The Euro Effect on Intra-EU Trade: Evidence from the Cross Sectionally Dependent Panel Gravity Models*

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Abstract

Recently, there has been an intense policy debate on the Euro effects on trade flows. The investigation of unobserved multilateral resistance terms in conjunction with omitted trade determinants has also assumed a prominent role in the literature. Following recent developments in panel data studies, we propose the cross-sectionally dependent panel gravity models. The desirable feature of this approach is to control for time-varying multilateral resistance, trade costs and globalisation trends through the use of both observed and unobserved factors, which are allowed to be cross-sectionally correlated. This approach also enables us to consistently estimate the impacts of (potentially endogenous) bilateral trade barriers. Applying the proposed approach to the dataset over 1960-2008 for 91 country-pairs of 14 EU countries, we find that the Euro impact on trade amounts to 3-4%, far less than those reported by earlier studies. Furthermore, the Euro is found to promote EU integration by eliminating exchange rate-related uncertainties. An obvious policy implication is that countries considering to join the Euro would benefit from the ongoing process of integration, but should also be wary of regarding promises of an imminent acceleration of intra-EU trade.

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Key Words: Gravity Models, Heterogeneous Panel Data, Cross-section Dependence, Multilateral Resistance, The Euro Effects on Trade and Integration.

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1 Introduction

It is generally acknowledged that the Euro-area economies become more integrated by adopting the Euro. Increased trade is one of the few undisputed gains from a currency union by eliminating exchange rate volatility and reducing the transactions costs of trade within the member countries. Thus, trade can be expected to rise. The important question is how much? The magnitude and the nature of the Euro's trade impact is not only important for the member countries, but also for the EU members who have not joined yet.

Gravity models of international trade have been widely employed to estimate the trading effect of common currency following the seminal paper by Rose (2000), where currency unions are found to increase trade by more than 200%. This huge effect has stimulated an intense debate in the literature, see Persson (2001), Alesina et al. (2002), Micco et al. (2003), Anderson and Van Wincoop (2004), Frankel (2005, 2008), Flam and Nordström (2006, 2007), Bun and Klaassen (2007), Berger and Nitsch (2008), de Nardis et al. (2008), Santos Silva and Tenreyro (2010), Herwartz and Weber (2010) and Camaero et al. (2012). Baldwin (2006) provides an extensive survey, establishing that the infamous Rose effect is severely (upward) biased. He then suggests that one should take into account time-varying country dummies as proxies for multilateral trade costs, and documents broad evidence that the currency-union effect can be greatly reduced by controlling for those components appropriately.

In a seminal paper, Anderson and van Wincoop (2003) propose the micro foundation of the gravity equation and emphasise an importance of taking into account multilateral resistance term (trade barriers of the pair of countries relative to those with respect to all trade partners). Acknowledging such an important issue, recently, the investigation of unobserved multilateral resistance terms in conjunction with omitted trade determinants has assumed a prominent role in measuring the Euro's trade effects more precisely (Baldwin, 2006; Baldwin and Taglioni, 2006). Bun and Klaassen (2007) claim that omitted trade determinants may cause the Euro effect to be substantially upward-biased, and suggest to introduce a time trend with heterogeneous coefficients, documenting that the Euro effect falls dramatically from 51% to 3%. Notice, however, (unobserved) multilateral resistance and trade costs are likely to capture history and time dependence of continuously deepening European integration in a rather complex manner. Such diverse measures might be better described by stochastic trending factors (e.g. Herwartz and Weber, 2010). This immediately raises another important issue of how best to model (unobserved) multilateral resistance and bilateral heterogeneity, which are likely to be cross-sectionally correlated. Some progress has been made. Recently, Behrens et al. (2012) derive a structural gravity equation, which allows both trade flows and error terms to be cross-sectionally correlated. They then propose the

modified spatial methodologies for the renewed analysis of infamous Canada–US border effects (McCallum, 1995). See also Camaero et al. (2012).

In this paper we follow recent developments in panel data studies (e.g. Ahn et al., 2001), and extend the cross-sectionally dependent panel gravity models advanced by Serlenga and Shin (2007) and Baltagi (2010). The desirable feature of this approach is to control for time-varying multilateral resistance, trade costs and globalisation trends through the use of observed and unobserved factors, which are explicitly modelled as (strong) cross-sectionally correlated (Chudik et al., 2011). Furthermore, this approach enables us to consistently estimate the impacts of (potentially endogenous) bilateral resistance such as the border effect through combining the estimators proposed by Pesaran (2006) and Bai (2009) with the instrument variables estimators advanced by Hausman and Taylor (1981), Amemiya and MaCurdy (1986), and Breusch, Mizon and Schmidt (1989). Within this framework, another important issue of the Euro effect on trade integration can be easily investigated by estimating the trend lines of time-varying coefficients of bilateral trade barriers.

We apply the proposed cross-sectionally dependent panel gravity model to the dataset over the period 1960-2008 (49 years) for 91 country-pairs of 14 EU member countries. Our main empirical findings are summarized as follows: Firstly, once we control for time-varying multilateral resistance terms and trade costs appropriately through cross-sectionally correlated unobserved factors, we find that the Euro impact on trade amounts to 3-4% only, which is also robust to the presence of trade diversion effects. Importantly, this magnitude is consistent with broad evidence compiled by Baldwin (2006) and more recent studies that attempt to address an importance of taking into account time-varying multilateral resistance and/or omitted trade determinants at least partially (e.g. Bun and Klaassen, 2007; Berger and Nitsch, 2008). We also find that the custom union effect is substantially reduced to 10% only. Combined together, these findings may support the thesis that the potential trade-creating effects of the Euro should be viewed in the proper historical and multilateral perspective rather than focusing simply on the formation of a monetary union as an isolated event.

Turning to the impacts of bilateral resistance terms, we find that the impacts of both distance and common language on trade are significantly negative and positive whereas the border impact is no longer significant. Further investigation of time-varying coefficients on these variables reveals that border and language effects started to decline more sharply just after 1999. The implication of these findings is that the Euro helps to reduce trade effects of bilateral resistance and thus promote integration among the Euro countries by eliminating exchange rate-related uncertainties and transaction costs. On the other hand, distance impacts have been rather stable, showing no pattern of downward trending. This generally

supports broad empirical evidence that the notion of the death of distance is difficult to identify in current trade data (Disdier and Head, 2008; Jacks, 2009).

The paper is organised as follows: Section 2 reviews recent literature on the Euro's Trade Effects. Section 3 describes the cross-sectionally dependent panel gravity models and proposes the associated estimation methodologies. Section 4 provides main empirical findings. Section 5 concludes.

2 Overview on the Euro's Trade Effects

Recently, there has been an intense policy debate on the Euro effects on trade flows between Euro and non-Euro nations.[1] Baldwin (2006) offers an extensive survey, covering the infamous Rose (2000)'s huge trading effect over 200% [2] as well as most recent studies reporting the relatively smaller effects. It is widely acknowledged that the Rose's estimate of the currency union effect on trade is severely (upward) biased. In particular, his estimates are heavily inflated by the presence of small (e.g. Ireland, Panama) or very small (e.g. Kiribati, Greenland, Mayotte) countries (Frankel, 2008). An important issue is why a currency union raises trades so much. In 2003 the UK Treasury made a bold prediction that the pro-trade effect of using the Euro on UK would be over 40%. One suspects that these results be seriously interpreted to mean that trade among its members would have collapsed in the late 1990s without the Euro (Santos Silva and Tenreyro, 2010). Thus, it is unclear whether one can uncover similar findings for the European monetary union involving the substantially large economies such as Germany and France.

The gravity model popularised by Rose (2000) attempts to provide the main link between trade flows and trade barriers, though his original approach has attracted the number of strong criticisms. The main critiques are classified as follows: inverse causality or endogeneity; missing or omitted variables; and incorrect model specification (nonlinearity or threshold effects). Nowadays, the general consensus is, once these methodological issues have been accommodated appropriately, that the currency union effect seems to be far less than those reported earlier by Rose and others, especially using the larger dataset including numerous smaller countries. Micco et al. (2003) provide the first evaluation of the Euro effect, and find that the common currency increases trade among Euro zone members by 4% in the short-run and 16% in the long-run. See also de Nardis and Vicarelli (2003), Flam

[1] Now, the euro area contains 17 EU member states. In 1999 eleven countries adopted the euro as a common currency while Greece entered in 2001. Slovenia joined in 2007, Cyprus and Malta in 2008, Slovakia in 2009 and Estonia in 2011. Denmark and the United Kingdom have 'opt-outs' from joining laid down in Protocols annexed to the Treaty whereas Sweden has not yet qualified to be part of the euro area.

[2] Rose (2000) estimates a gravity equation using data for 186 countries from 1970 to 1990 and finds that countries in a currency union trade three times as much.

and Nordström (2006), Berger and Nitsch (2008), and de Nardis et al. (2008), from which we still find that the range of the estimated Euro effects is very wide from 2% to more than 70%.

Frankel (2005) claims that there are other third factors, such as common language, colonial history, and political/institutional link, that may influence both currency choice and trade link. In this regard, high correlations reported in earlier studies may be spurious as an artifact of reverse causality. A related issue is how the currency union is formed. Countries who decide to join a currency union are self-selected on the basis of distinctive features shared by countries that have been EU members during the pre-Euro period. Hence, countries are likely to foster integration by enhancing standards of harmonization and reducing regulatory barriers. To address this issue, a number of studies have employed different techniques such as Heckman selection and instrumental variables, though they still obtained the substantial Euro effects on trade, e.g. Persson (2001) and Alesina et al. (2002).[3]

A more important issue is omitted variables bias. Omitted pro-bilateral trade variables are likely to be correlated with the currency union dummy, as the formation of currency unions is not random, but rather driven by some factors which are likely to be omitted from the gravity regression. The implication is that the Euro effect will capture general economic integration among the member states, not merely the currency effect. Several studies tried to reduce the endogenous effect of currency unions by introducing country-pair and year fixed effects in the gravity regression, see Micco et al. (2003), Flam and Nordström (2006) and Berger and Nitsch (2008).

Anderson and van Wincoop (2003) propose the 'micro foundation' of the gravity equation by introducing the multilateral resistance terms, which are relative trade barriers -the bilateral barrier relative to average trade barriers that both countries face with all their trading partners. The empirical gravity literature has simply added so-called remoteness variables, which are defined as a weighted average distance from all trading partners with the weights being based on the size of the trading partners, e.g. Frankel and Wei (1998) and Melitz (2007), though such atheoretical remoteness indices fail to capture any of the relative trade barriers in a coherent manner. Hence, the standard gravity model is seriously lacking if multilateral resistance terms and/or trade costs are ignored or seriously misspecified.

Furthermore, Baldwin (2006) stresses an importance of taking into account time-varying multilateral resistance terms, and criticises the conventional fixed effect estimation technique because many of omitted pair-specific variables clearly reflect time-varying factors such as multilateral trade costs (Anderson and von

[3] The Heckman approach produces estimates in the order of 50 %. Surprisingly, however, the instrumental variable approach generates huge estimates of the currency effects, sometimes even larger than the Rose effect.

Wincoop, 2004). The use of time-invariant effects only may still leave a times-series trace in the residual, which is likely to be correlated with the currency union dummy. Baldwin and Taglioni (2006) claim that an incomplete account of time variation in multilateral resistance terms is likely to cause omitted-variable bias. Therefore, once such time-varying effects are explicitly incorporated in the gravity model, the impact of currency-union would be greatly reduced. A number of studies have attempted to capture time-varying effects when estimating the Euro's trade effect. In particular, Bun and Klaassen (2007) claim that upward trends in omitted trade determinants may cause the estimated Euro effect to be substantially upward-biased, and these biases tend to be magnified as the sample period enlarges. In order to deal with different effects of time varying omitted components across country-pairs, they introduce a time trend with heterogeneous coefficients, and find that the Euro effect on trade falls dramatically from 51% to 3% for the dataset over the period, 1967-2002. Moreover, Berger and Nitsch (2008) find no impact of the Euro on trade when including a linear trend in the gravity regression with the data from 1948 to 2003, and conclude that the Euro-12 countries have already been integrated strongly even before the Euro was created.

In sum, a large number of existing studies have established the importance of appropriately taking into account unobserved and time-varying multilateral resistance and bilateral heterogeneity, simultaneously. This immediately raises another important issue of cross-section dependence among trade flows, which has been so far neglected. Only recently, Herwartz and Weber (2010) propose to capture both multilateral resistance terms and omitted trade costs via unobserved time-varying country-pair specific random walk factors, and develop the Kalman-filter extension of the gravity model. They find that aggregate trade (export) within the Euro area increases between 2000 and 2002 by 15 to 25 percent compared with trade with non-members of the Euro area due to a decrease in long-lasting trade costs. Camaero et al. (2012) suggest to estimate a gravity equation by a panel-based cointegration approach that allows for cross-section dependence through the common factors. Applying the continuously updated estimator of Bai et al. (2009) to the bilateral dataset for 26 OECD countries during the period 1967-2008, they find that the Euro appears to generate somewhat lower trade effects than suggested by previous studies.[4]

More importantly, Behrens et al. (2012) derive a quantity-based structural gravity equation system in which both trade flows and error terms are allowed to be cross-sectionally correlated, and propose the modified spatial techniques by adopting a broader definition of the spatial weight matrix, called the interaction

[4] The approach by Camaero et al. (2012) can be regarded as an extension of Bun and Klaasen (2007), who estimate the long-run cointegrating relationship without controlling for cross-section dependence. Interestingly, however, the euro impact is estimated at 16-7%, substantially higher than 3% estimated by Bun and Klaasen (2007).

matrix, which can be derived directly from theoretical model. By controlling for cross-sectional interdependence and thus directly capturing multilateral resistance, they find that the measured Canada–US border effects are significantly lower than paradoxically large estimates reported by McCallum (1995). Behrens et al. (2012) also argue that their approach - unconstrained linearized gravity equation with cross-sectionally correlated trade flows - is better suited than the two-stage gravity equation system with nonlinear constraints in unobservable price indices advanced by Anderson and von Wincoop (2003).[5]

Taken together, all of the above discussions may suggest that an Euro effect on trade is expected to be smaller in the future than previously thought once multilateral resistance term is well-captured via the cross-sectional interdependence of trade flows. In retrospect, Serlenga and Shin (2007, henceforth SS) is the first paper to introduce the cross-section dependence into the panel gravity model, and to provide consistent estimation procedure for both time-varying and time-invariant regressors. Employing the dataset over the period 1960-2001, SS find that the introduction of a common currency does not exert any significant effect on intra-EU trade, though their sample covers only three years' data since the introduction of the Euro in 1999. Given the availability of a longer sample, we wish to redress this important issue by extending the cross-sectionally dependent panel gravity model and addressing all of the issues related to unobserved and time-varying multilateral resistance and bilateral heterogeneity as surveyed above.

3 Panel Gravity Models in the Presence of Cross Section Dependence

The fixed effect estimation has been most popular in the literature on gravity models, e.g. Rose and van Wincoop (2001), though it fails to estimate coefficients on time-invariant variables such as distance, since the within transformation wipes out those variables. Another important issue is how to extend the fixed effect specification into a more general case with individual-specific and time-varying effects, both of which affect bilateral trade flows. The multilateral resistance function and trade costs are not only difficult to measure, but also likely to vary over time. A number of approaches have been proposed. Simply, fixed time dummies or time trends are added as a proxy for time-varying effects in the gravity equation, e.g. Baldwin and Taglioni (2006), and Berger and Nitsch (2008). Bun and Klaassen (2007) allow to time trend coefficients to be heterogeneous across country-pairs. Alternatively, some studies include *ad hoc* regional remoteness indices (e.g. Melitz,

[5] Anderson and van Wincoop (2003) obtain the multilateral resistance terms in the first stage, and use them in the second stage gravity equation using nonlinear least squares.

2007), although these indices have no theoretical foundation (Behrens et al., 2012).

We now consider a more generalized panel data model advanced by SS and Baltagi (2010):

$$y_{it} = \boldsymbol{\beta}'\mathbf{x}_{it} + \boldsymbol{\gamma}'\mathbf{z}_i + \boldsymbol{\pi}_i'\mathbf{s}_t + \varepsilon_{it},\ i = 1, ..., N,\ t = 1, ..., T, \tag{1}$$

$$\varepsilon_{it} = \alpha_i + \boldsymbol{\varphi}_i'\boldsymbol{\theta}_t + u_{it}, \tag{2}$$

where $\mathbf{x}_{it} = (x_{1,it}, ..., x_{k,it})'$ is a $k \times 1$ vector of variables that vary across individuals and over time periods, $\mathbf{s}_t = (s_{1,t}, ..., s_{s,t})'$ is an $s \times 1$ vector of observed time-specific factors, $\mathbf{z}_i = (z_{1,i}, ..., z_{g,i})'$ is a $g \times 1$ vector of individual-specific variables, $\boldsymbol{\beta} = (\beta_1, ..., \beta_k)'$, $\boldsymbol{\gamma} = \left(\gamma_1, ..., \gamma_g\right)'$ and $\boldsymbol{\pi}_i = (\pi_{1,i}, ..., \pi_{s,i})'$ are conformably defined column vectors of parameters, α_i is an individual effect that might be correlated with explanatory variables $\mathbf{x}_{it}$ and $\mathbf{z}_i$, $\boldsymbol{\theta}_t$ is the $c \times 1$ vector of unobserved common factors with conformable parameter vector, $\boldsymbol{\varphi}_i = \left(\varphi_{1,i}, ..., \varphi_{c,i}\right)'$, and u_{it} is a zero mean idiosyncratic uncorrelated random disturbance.

The distinguishing feature of this model is that it allows for both observed and unobserved time effects both of which are cross-sectionally correlated. Both factors are expected to provide good proxy for any remaining complex time-varying patterns associated with multilateral resistance and globalisation trends, e.g. Mastromarco et al. (2012). Notice that the cross-section dependence in (1) is explicitly allowed through heterogeneous loadings, $\boldsymbol{\varphi}_i$, see Pesaran (2006) and Bai (2009).[6] It is easily seen that most specifications of the gravity equation in the literature can be expressed as a variation of the model given by (1) and (2).[7] Hence, this factor-based cross-sectionally dependent panel gravity model is expected to capture the time-varying pattern of unobserved trading effects, such as the multilateral resistance, in a robust manner.

The conventional panel data estimators of (1) would be seriously biased without properly accommodating the cross-sectionally dependent factor structure given by (2). To explicitly address this issue, we consider the two leading approaches proposed by Pesaran (2006) and Bai (2009). Following Pesaran (2006) we first derive the cross-sectionally augmented regression of (1) as follows:

$$y_{it} = \boldsymbol{\beta}'\mathbf{x}_{it} + \boldsymbol{\gamma}'\mathbf{z}_i + \boldsymbol{\lambda}_i'\mathbf{f}_t + \alpha_i^* + u_{it}^*,\ i = 1, ..., N,\ t = 1, ..., T, \tag{3}$$

where $\mathbf{f}_t = (\mathbf{s}_t', \bar{y}_t, \bar{\mathbf{x}}_t')' \left\{= (f_{1,t}, ..., f_{\ell,t})'\right\}$ is the $\ell \times 1$ vector of augmented time-specific factors with $\ell = s + 1 + k$ and $\boldsymbol{\lambda}_i = (\lambda_{1,i}, ..., \lambda_{\ell,i})'$, $\bar{y}_t = N^{-1}\sum_{i=1}^{N} y_{it}$, $\bar{\mathbf{x}}_t =$

[6] Chudik et al. (2011) show that these factor models exhibit the strong form of cross-sectional dependence since the maximum eigenvalue of the covariance matrix for ε_{it} tends to infinity at rate N. On the other hand spatial econometric models, developed by Behrens et al. (2012), display the weak form of cross-sectional dependence, which can be represented by an infinite number of weak factors and no idiosyncratic error.

[7] For example, the specification employed by Bun and Klaassen (2007) is obtained by simply including t as one element in $\mathbf{s}_t$, but without unobserved factors, $\boldsymbol{\theta}_t$.

$N^{-1}\sum_{i=1}^{N}\mathbf{x}_{it}$, $\boldsymbol{\lambda}_i' = (\boldsymbol{\pi}_i' - (\varphi_i/\bar{\varphi})\,\bar{\boldsymbol{\pi}}', (\varphi_i/\bar{\varphi}), -(\varphi_i/\bar{\varphi})\,\boldsymbol{\beta}')'$ with $\bar{\varphi} = N^{-1}\sum_{i=1}^{N}\varphi_i$ and $\bar{\boldsymbol{\pi}} = N^{-1}\sum_{i=1}^{N}\boldsymbol{\pi}_i$, $\alpha_i^* = \alpha_i - (\varphi_i/\bar{\varphi})\,\bar{\alpha} - (\varphi_i/\bar{\varphi})\,\boldsymbol{\gamma}'\bar{\mathbf{z}}$ with $\bar{\alpha} = N^{-1}\sum_{i=1}^{N}\alpha_i$ and $\bar{\mathbf{z}} = N^{-1}\sum_{i=1}^{N}\mathbf{z}_i$, and $u_{it}^* = u_{it} - (\varphi_i/\bar{\varphi})\,\bar{u}_t$ with $\bar{u}_t = N^{-1}\sum_{i=1}^{N}u_{it}$. Using (3), we can derive the CCEP estimator of $\boldsymbol{\beta}$ in (4) by (4) below. Alternatively, $\boldsymbol{\beta}$ can be consistently estimated by the principal component (PC) estimator proposed by Bai (2009). Chudik et al. (2011) show that the PC estimator is similar to the CCEP estimator, except that the cross section averages are replaced by the estimated common factors $\left(\hat{\boldsymbol{\theta}}_t\right)$, which are obtained using the Bai and Ng (2002) procedure.[8] In this case we have $\mathbf{f}_t = \left(\mathbf{s}_t', \hat{\boldsymbol{\theta}}_t'\right)'$ in (3). Specifically, the CSD-consistent estimator of $\boldsymbol{\beta}$ is given by[9]

$$\hat{\boldsymbol{\beta}}_{CSD} = \left(\sum_{i=1}^{N}\mathbf{x}_i'\mathbf{M}_T\mathbf{x}_i\right)^{-1}\left(\sum_{i=1}^{N}\mathbf{x}_i'\mathbf{M}_T\mathbf{y}_i\right),\ \hat{\boldsymbol{\beta}}_{CSD} = \hat{\boldsymbol{\beta}}_{CCEP}\ or\ \hat{\boldsymbol{\beta}}_{PC} \tag{4}$$

where $\mathbf{y}_i = (y_{i1}, ..., y_{iT})'$, $\mathbf{x}_i = (\mathbf{x}_{i1}, ..., \mathbf{x}_{iT})'$, $\mathbf{M}_T = \mathbf{I}_T - \mathbf{H}_T\left(\mathbf{H}_T'\mathbf{H}_T\right)^{-1}\mathbf{H}_T'$, $\mathbf{H}_T = (\mathbf{1}_T, \mathbf{f})$, $\mathbf{1}_T = (1, ..., 1)'$ and $\mathbf{f} = (\mathbf{f}_1', ..., \mathbf{f}_T')'$.

Both CCEP and PC estimators are unable to estimate the coefficients, $\boldsymbol{\gamma}$ on time-invariant variables because they are the extended fixed effect estimators. In this regard, SS combine the CCEP estimation with the instrumental variables estimation proposed by Hausman and Taylor (1981, HT), and develop the CCEP-HT estimator. Baltagi (2010) further proposes the CCEP-AM estimator by employing the additional instrument variables proposed by Amemiya and MaCurdy (1986, AM). It is then straightforward to consider further additional set of instrument variables proposed by Breusch, Mizon and Schmidt (1989, BMS), which we denote by the CCEP-BMS estimator. We can also develop the corresponding counterparts, using the Bai's PC estimator, which we denote by PC-HT, PC-AM and PC-BMS estimators, respectively.

Conformable with HT, we decompose $\mathbf{x}_{it} = (\mathbf{x}_{1it}', \mathbf{x}_{2it}')'$ and $\mathbf{z}_i = (\mathbf{z}_{1i}', \mathbf{z}_{2i}')'$, where $\mathbf{x}_{1it}$, $\mathbf{x}_{2it}$ are $k_1 \times 1$ and $k_2 \times 1$ vectors, and $\mathbf{z}_{1i}$, $\mathbf{z}_{2i}$ are $g_1 \times 1$ and $g_2 \times 1$ vectors. Then, we can estimate $\boldsymbol{\gamma}$ consistently using instrumental variables in the following regression:

$$d_{it} = \boldsymbol{\gamma}_1'\mathbf{z}_{1i} + \boldsymbol{\gamma}_2'\mathbf{z}_{2i} + \alpha_i^* + u_{it}^* = \mu + \boldsymbol{\gamma}'\mathbf{z}_i + \varepsilon_{it}^*,\ i = 1, ..., N,\ t = 1, ..., T, \tag{5}$$

where $d_{it} = y_{it} - \hat{\boldsymbol{\beta}}_{CSD}'\mathbf{x}_{it} - \boldsymbol{\lambda}_i'\mathbf{f}_t$, $\mu = E(\alpha_i^*)$ and $\varepsilon_{it}^* = (\alpha_i^* - \mu) + u_{it}^*$ is a zero mean process by construction. In matrix notation we have:

$$\mathbf{d} = \mu\mathbf{1}_{NT} + \mathbf{Z}_1\boldsymbol{\gamma}_1 + \mathbf{Z}_2\boldsymbol{\gamma}_2 + \boldsymbol{\varepsilon}^*, \tag{6}$$

[8] After applying the within transformation of the model, (1), we can extract the factors from the within residuals in an iterative manner.

[9] Under fairly standard regularity conditions, Pesaran (2006) and Bai (2009) prove that as $(N, T) \to \infty$ jointly, $\hat{\boldsymbol{\beta}}_{CSD}$ is consistent and follows the asymptotic normal distribution.

where $\mathbf{d} = (\mathbf{d}_1', ..., \mathbf{d}_N')'$, $\mathbf{d}_i = (d_{i1}, ..., d_{iT})'$, $\mathbf{Z}_j = \left(\left(\mathbf{z}_{j1}' \otimes \mathbf{1}_T\right)', ..., \left(\mathbf{z}_{jN}' \otimes \mathbf{1}_T\right)'\right)'$, $j = 1, 2$, $\mathbf{1}_{NT} = (\mathbf{1}_T', ..., \mathbf{1}_T')'$, $\mathbf{1}_T = (1, ..., 1)'$, and $\boldsymbol{\varepsilon}^* = \left(\boldsymbol{\varepsilon}_1^{*\prime}, ..., \boldsymbol{\varepsilon}_N^{*\prime}\right)'$ with $\boldsymbol{\varepsilon}_i^* = (\varepsilon_{i1}^*, ..., \varepsilon_{iT}^*)'$. Replacing $\mathbf{d}$ by its consistent estimate, $\hat{\mathbf{d}} = \left\{\hat{d}_{it}, i = 1, ..., N,\ t = 1, ..., T,\right\}$, where $\hat{d}_{it} = y_{it} - \hat{\boldsymbol{\beta}}_{CSD}'\mathbf{x}_{it} - \hat{\boldsymbol{\lambda}}_i'\mathbf{f}_t$ and $\hat{\boldsymbol{\lambda}}_i$ are the OLS estimators of $\boldsymbol{\lambda}_i$ consistently estimated from the regression of $\left(y_{it} - \hat{\boldsymbol{\beta}}_{CSD}'\mathbf{x}_{it}\right)$ on $(1, \mathbf{f}_t)$ for $i = 1, ..., N$, we now have:

$$\hat{\mathbf{d}} = \mu\mathbf{1}_{NT} + \mathbf{Z}_1\boldsymbol{\gamma}_1 + \mathbf{Z}_2\boldsymbol{\gamma}_2 + \boldsymbol{\varepsilon}^+ = \mathbf{C}\boldsymbol{\delta} + \boldsymbol{\varepsilon}^+, \tag{7}$$

where $\boldsymbol{\varepsilon}^+ = \boldsymbol{\varepsilon}^* + \left(\hat{\mathbf{d}} - \mathbf{d}\right)$, $\mathbf{C} = (\mathbf{1}_{NT}, \mathbf{Z}_1, \mathbf{Z}_2)$ and $\boldsymbol{\delta} = (\mu, \boldsymbol{\gamma}_1', \boldsymbol{\gamma}_2')'$.

To deal with nonzero correlation between $\mathbf{Z}_2$ and $\boldsymbol{\alpha}$, we need to find the $NT \times (1 + g_1 + h)$ matrix of instrument variables:

$$\mathbf{W} = [\mathbf{1}_{NT}, \mathbf{Z}_1, \mathbf{W}_2],$$

where $\mathbf{W}_2$ is an $NT \times h$ matrix of instrument variables for $\mathbf{Z}_2$ with $h \geq g_2$ for identification. First, we follow SS and consider the $NT \times (k_1 + \ell)$ HT instrument matrix given by

$$\mathbf{W}_2^{HT} = \left[\mathbf{P}\mathbf{X}_1, \mathbf{P}\hat{\boldsymbol{\xi}}_1, \mathbf{P}\hat{\boldsymbol{\xi}}_2, ..., \mathbf{P}\hat{\boldsymbol{\xi}}_\ell\right]$$

where $\mathbf{P} = \mathbf{D}(\mathbf{D}'\mathbf{D})^{-1}\mathbf{D}'$ is the $NT \times NT$ idempotent matrix with $\mathbf{D} = \mathbf{I}_N \otimes \mathbf{1}_T$, $\mathbf{I}_N$ being an $N \times N$ identity matrix, and $\hat{\boldsymbol{\xi}}_j = \left(\hat{\lambda}_{j,1}\mathbf{f}_j', \hat{\lambda}_{j,2}\mathbf{f}_j', ..., \hat{\lambda}_{j,N}\mathbf{f}_j'\right)'$, $j = 1, ..., \ell$, where $\mathbf{f}_j = (f_{j,1}, ..., f_{j,T})'$ with $\hat{\lambda}_{j,i}$ being consistent estimate of heterogenous factor loading, $\lambda_{j,i}$. Next, we follow Baltagi (2010) and derive the $NT \times (k_1 + \ell + Tk_1 + T\ell)$ AM instrument matrix by

$$\mathbf{W}_2^{AM} = \left[\mathbf{W}_2^{HT}, (\mathbf{Q}\mathbf{X}_1)^*, \left(\mathbf{Q}\hat{\boldsymbol{\xi}}_1\right)^*, \left(\mathbf{Q}\hat{\boldsymbol{\xi}}_2\right)^*, ..., \left(\mathbf{Q}\hat{\boldsymbol{\xi}}_\ell\right)^*\right] \tag{8}$$

where $\mathbf{Q} = \mathbf{I}_{NT} - \mathbf{P}$ and $(\mathbf{Q}\mathbf{X}_1)^* = (\mathbf{Q}\mathbf{X}_{11}, \mathbf{Q}\mathbf{X}_{12}, ..., \mathbf{Q}\mathbf{X}_{1T})$ is the $NT \times k_1T$ matrix with $\mathbf{Q}\mathbf{X}_{1t} = (\mathbf{Q}\mathbf{X}_{11t}, ..., \mathbf{Q}\mathbf{X}_{1kt})'$.[10] Finally, it is straightforward to derive the $NT \times (k_1 + \ell + Tk_1 + T\ell + Tk_2)$ BMS instrument matrix by

$$\mathbf{W}_2^{BMS} = \left[\mathbf{W}_2^{AM}, (\mathbf{Q}\mathbf{X}_2)^*\right]$$

where $(\mathbf{Q}\mathbf{X}_2)^* = (\mathbf{Q}\mathbf{X}_{21}, \mathbf{Q}\mathbf{X}_{12}, ..., \mathbf{Q}\mathbf{X}_{2T})$.[11]

To derive the consistent estimator of $\boldsymbol{\delta}$, we premultiply $\mathbf{W}'$ by (7)

$$\mathbf{W}'\hat{\mathbf{d}} = \mathbf{W}'\mathbf{C}\boldsymbol{\delta} + \mathbf{W}'\boldsymbol{\varepsilon}^+. \tag{9}$$

[10] Notice that the rank of $(\mathbf{Q}\mathbf{X}_1)^*$ is $(T-1)k_1$, because only $(T-1)$ deviations from means are (linearly) independent since each variable (see BMS). Similarly for $\left(\mathbf{Q}\hat{\boldsymbol{\xi}}_1\right)^*$, ..., $\left(\mathbf{Q}\hat{\boldsymbol{\xi}}_\ell\right)^*$.

[11] As before, the rank of $(\mathbf{Q}\mathbf{X}_2)^*$ is only $(T-1)k_2$.

Therefore, the GLS estimator of $\boldsymbol{\delta}$ is obtained by

$$\hat{\boldsymbol{\delta}}_{GLS} = \left[\mathbf{C}'\mathbf{W}\mathbf{V}^{-1}\mathbf{W}'\mathbf{C}\right]^{-1}\mathbf{C}'\mathbf{W}\mathbf{V}^{-1}\mathbf{W}'\hat{\mathbf{d}}, \tag{10}$$

where $\mathbf{V} = Var\left(\mathbf{W}'\boldsymbol{\varepsilon}^{+}\right)$. To obtain the feasible GLS estimator we replace $\mathbf{V}$ by its consistent estimator. In practice, estimates of $\boldsymbol{\delta}$ and $\mathbf{V}$ can be obtained iteratively until convergence, see also SS for further details.

Notice that the HT-IV estimator employs only the mean of $\mathbf{X}_1$ to be uncorrelated with the effects, α_i^* whereas the AM-IV estimator exploits such moment conditions to be held at every time period. Hence, the validity of the AM instruments requires a stronger exogeneity assumption for $\mathbf{X}_1$, under which the AM-IV estimator is more efficient than HT-IV. Furthermore, the BMS instruments require uncorrelatedness of $\mathbf{X}_2$ with α_i^* separately at every point in time. The validity of the AM and the BMS instruments can be easily tested via the Hausman statistics testing for the difference between HT-IV and AM-IV estimators and between AM-IV and BMS-IV estimators as follows:

$$\begin{aligned} H_{AM} &= \left(\hat{\boldsymbol{\delta}}_{AM} - \hat{\boldsymbol{\delta}}_{HT}\right)' \left[Var\left(\hat{\boldsymbol{\delta}}_{HT}\right) - Var\left(\hat{\boldsymbol{\delta}}_{AM}\right)\right]^{-1} \left(\hat{\boldsymbol{\delta}}_{AM} - \hat{\boldsymbol{\delta}}_{HT}\right) \\ H_{BMS} &= \left(\hat{\boldsymbol{\delta}}_{BMS} - \hat{\boldsymbol{\delta}}_{AM}\right)' \left[Var\left(\hat{\boldsymbol{\delta}}_{AM}\right) - Var\left(\hat{\boldsymbol{\delta}}_{BMS}\right)\right]^{-1} \left(\hat{\boldsymbol{\delta}}_{BMS} - \hat{\boldsymbol{\delta}}_{AM}\right) \end{aligned}$$

both of which follow the asymptotic χ_g^2 distribution with the degree of freedom g being the number of coefficients tested.

4 Empirical Results

We extend the dataset analysed by Serlenga and Shin (2007) to cover the longer period 1960-2008 (49 years) for 91 country-pairs amongst 14 EU member countries (Austria, Belgium-Luxemburg, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal, Spain, Sweden, United Kingdom).[12] Our sample period consists of several important economic integrations, such as the Custom Union in 1958, the European Monetary System in 1979 and the Single Market in 1993, all of which can be regarded as promoting intra-EU trades (Eurostat, 2008).[13] Given that the Euro effect should be analysed as an ongoing process (Berger and Nitsch, 2008), we wish to examine the Euro's trading effect

[12] Denmark, Sweden and The UK constitute a meaningful control group since these nonmember countries, as part of the EU, experienced similar history and faced similar legislation and regulation to euro-area countries.

[13] To mitigate the potentially negative impacts of the ongoing global financial crisis on our analysis, however, we exclude the data after 2008. Both imports and exports in the Euro area fell by around one-fifth in 2009 (Statistical Yearbook, 2010).

more precisely by applying the extended cross-sectionally dependent panel data methodology developed in Section 3 to the dataset with the longer sample period.[14]

Focussing on the EU trade patterns since the Euro, we find it interesting to observe from Eurostat (2003) that the EU trade fell by only 0.7% per annum during 2000-2003, even though the global trades sharply contracted following the world-wide recession in 2001 and 2002 (trade flows of US, Japan and Canada, recorded an annual reduction of around 6.7%). The EU trades grew strongly during 2003-2007, thanks to upswing in the world trade taking place after 2003 and the accession of 12 new member states in 2004 and 2007. In particular, the intra-EU trade increased by almost 40% during 2003-2004, mainly due to the 25% real appreciation of the Euro against the US dollar (Eurostat, 2003). The Euro area (intra and extra) trade in goods grew significantly over the last decade - increased to 32% of Euro area GDP in 2008 from 26 % in 1999 (Unctad, 2012). Furthermore, trade growth was faster than real GDP growth, leading to an increasing openness ratio of the Euro area (as measured by the sum of imports and exports as a share of GDP in real terms), which reached 82% in 2008 as compared to 64% in 1999 (World Bank, 2012). These tight trade linkages can be explained partially by the existence of both single market and single currency (ECB Bulletin, 2010).

Table 1 presents key summary figures of EU trade shares and growths.[15] First, the intra-EU trade has been a considerable part of the total trade in EU. Its share reached and has stayed over 60% since 1990s. Second, the US is still the leading trade partner of the EU, though its leading role has recently been challenged by China and Russia, as the US share of extra-EU trade decreased significantly from 21.9% in 2000 to 15.1% in 2008. Third, the trade still grows faster than real GDP in 2000s. Finally, the share of exports is slightly higher than that of imports.

Table 1 about here

When estimating the panel data model of gravity, (1) and (2), we consider two cases. In the first case, we consider the basic model without unobserved time-varying factors, $\varphi_i \theta_t$ in (2).[16] Here, we consider two subcases: first, the model

[14] The dependent variable is the logarithm of real total trade. The regressors are the logarithm of total GDP ($TGDP$) which proxies for trade partners' mass; similarity in size (SIM) and difference in relative factor endowment (RLF) which are introduced following recent advancements of New Trade Theory; the logarithm of real exchange rate (RER) proxying for relative price effects; a dummy for European Community membership (CEE) and a dummy for European Monetary Union (EMU); time-invariant bilateral resistance terms such as a dummy for common language (LAN), a dummy for common border (BOR), and the logarithm of geographical distance (DIS). See the Data Appendix in SS for details of the data construction.

[15] This is the updated table as reported in Serlenga and Shin (2007).

[16] Notice that the estimation results for the basic models are presented mainly for comparison with existing studies.

without any observed factors; namely, $s_t = \{\varnothing\}$. Secondly, we include linear time trends as an observed factor in (1); namely, $s_t = \{t\}$. Next, we consider our proposed model which explicitly incorporates unobserved time-varying factors, $\varphi_i \theta_t$ in (2).[17] Here, we also consider two subcases with and without linear time trends as an observed factor. In each case, in order to consistently estimate all of the parameters in the model, we employ two alternative estimators, the CCEP- and the PC-based estimators, as described in details in Section 3.

Table 2 presents the estimation results for the basic model without cross-section dependence and with individual effects only, using the alternative estimation methodologies mostly employed in the empirical literature of gravity models.[18] The REM assumption that there is no correlation between regressors and individual effects is convincingly rejected in all cases considered. Therefore, we focus on the FEM results. The FEM estimation results are all statistically significant and consistent mostly with our *a priori* expectations. The impact of $TGDP$ (the sum of home and foreign country GDPs) on trade is positive. Similarity in size (SIM) significantly boosts trade flows while the impact of relative difference in factor endowments between trading partners (RLF) is small but significantly positive. A depreciation of the home currency (increase in RER) leads to an increase in trade flows, reflecting the fact that the export component of the total trade is larger than the import component (see Table 1). Trade and currency union memberships (CEE and EMU) significantly boost real trade flows, although the magnitudes of both impacts (0.31 and 0.14) seem to be somewhat too high, confirming our main concern that upward trends in omitted trade determinants may cause these effects to be substantially upward-biased.

We now turn to estimating the impacts of individual-specific variables, jointly. Under the maintained assumption that LAN is the only time invariant variable correlated with individual effects (as common language is a proxy for cultural and historical proximity), we select the final set of instruments containing RER and RLF, after conducting a sequence of the Sargan tests for checking the validity of over-identifying restrictions.[19] The HT-IV and AM-IV estimation results reported

[17] We have also estimated the model with conventional two-way fixed effects, which fails to accommodate cross-section dependence. We find that the estimation results - available upon request - are mostly misleading, as also strongly highlighted by SS.

[18] The pooled OLS (POLS) is likely to gain in efficiency but biased due to neglected (individual) heterogeneity. The fixed effects model (FEM) explicitly takes into account the bilateral trade heterogeneity. The random effects model (REM) would estimate the parameters on both time-varying and time-invariant variables potentially more efficiently.

[19] In principle, for the AM case, we can create T instruments for each exogenous variable in $\mathbf{X}_1$, and obtain the set of Tk_1 instruments. But, this is valid only if the variables in $\mathbf{X}_1$ vary over i and t (BMS). If any variable in X_1 varies only over t but is constant across i - only one of T possible instruments can be used. Also, if any variable in $\mathbf{X}_1$ shows (very) low variation across i and over t, the $NT \times T$ matrix $(\mathbf{QX}_1)^*$ may not have the full rank. In this case only

in columns 4 and 5 of Table 2, show that the impacts of distance (DIS) and common language (LAN) on trade are significantly negative and positive, respectively. On the other hand, the common border impact is statistically insignificant (the signs of HT-IV and AM-IV are negative and positive). When comparing the IV estimation results to the (biased) POLS counterparts, we find that the impact of distance generally decreases (in absolute value) while that of common language is higher, especially for HT-IV. Because the Hausman test based on the contrast between HT-IV and AM-IV estimates does not reject the legitimacy of the AM-IV estimates, we focus on more efficient AM-IV results, and find that impacts of DIS and LAN are significant (-0.61 versus 0.87). On the other hand, the border impact is still insignificant and thus negligible.

When comparing our current estimation results with those reported earlier by SS employing the data with the shorter sample periods (1960-2001), we find that both estimation results are more or less similar, though there are two important differences to notice: First, the border effect is now no longer significant at all. This evidence suggest that geographic contiguity is no longer adequate to measure economic or transportation costs given recent improvements of information and communication technologies (e.g. Brun *et al.* 2005; Melitz, 2007). Secondly, the Euro impact on trade is statistically significant and almost doubles (rising to 0.14 from 0.08 with the earlier shorter samples). According to the recent survey evidence discussed in Section 2, however, these (inflated) effects are likely to capture neglected upward trends in omitted trade determinants.

Table 2 about here

Table 3 presents the estimation results for the basic gravity model with linear trends as an observed factor, namely, $\mathbf{s}_t = t$ in (1). Notice that estimation results in Table 3 are quite similar to those reported in Table 2. So focussing on trade and currency union effects on trade, we find that the Euro effect falls somewhat from 0.14 to 0.11 while the coefficients on CEE are almost the same. Notice that this is similar to an empirical specification employed by Bun and Klaassen (2007), in order to address the issue of upward-biased euro effects due to omitted trade determinants, who document that the Euro effect on trade falls dramatically from 51% to 3%. We find that the cross-section average of heterogeneous country-pair trend coefficients (from the FEM) is positive (0.18) and significant at 1% level, providing a partial support for the use of heterogeneous time trends as a proxy for capturing upward time-varying omitted effects. As our findings suggest, however, only the introduction of heterogeneous time trends is not yet sufficiently effective

a subset of the T instruments can be used. We select the final subset on the basis of the Sargan test results. Indeed, this caveat applies to RER_{it}, which shows (very) low variation across i and over t after 1999.

in capturing any upward trends in omitted trade determinants even after we have controlled for several trade determinants including exchange rates (relative price effects), similarity and relative difference in factor endowments. Claiming that European integration is continuously deepening, Berger and Nitsch (2008) also suggest to approximate such integration by a deterministic time trend. Given that (unobserved) multilateral resistance terms and trade costs are likely to exhibit a certain degree of history and time dependence in a complex manner, however, such diverse measures might be better described by stochastic trending factors (e.g. Herwartz and Weber, 2010). Hence, to address these important issues, we turn to our proposed models.

Table 3 about here

Table 4 displays the estimation results for the extended model with cross-section dependence and without linear trends as an observed factor. This is the factor-augmented panel data model, which explicitly takes into account cross section dependence. Here, we consider two consistent estimators, CCEP and PC estimators. In the former case we consider $\mathbf{f}_t = \left\{RERT_t, \overline{TGDP}_t, \overline{SIM}_t, \overline{RLF}_t, \overline{CEE}_t\right\}'$ in (3), where the bar over variables indicates their cross-sectional average, and $RERT_t$ is an observed factor defined as the (logarithm of) real exchange rates that would capture relative price effects between the European currencies and the US dollar. To derive PC estimators, we first extract six common PC factors using the Bai and Ng (2002) procedure, and use them as $\mathbf{f}_t$ in (3) together with $RERT_t$.[20] Furthermore, in order to consistently estimate the impacts of individual-specific variables jointly under the maintained assumption that LAN is the only time invariant variable correlated with individual effects, we use the same instrument variables, $IV = \{RER_{it}, RLF_{it}\}$. We also consider an additional instrument set, denoted $IV1 = \left\{IV, \hat{\boldsymbol{\xi}}_{it}\right\}$, where $\hat{\xi}_{it} = \hat{\lambda}_i f_t$, and $\hat{\lambda}_i$ are estimated loadings.[21]

We now find the following main differences between the estimation results summarised in Tables 2 and 4: First, the impact of RLF is no longer significant,

[20] After estimating a number of specifications augmented with several combinations of factors, we have selected the final specification on the basis of overall statistical significance and empirical coherence. Overall results are qualitatively similar and robust to different factor specifications.

[21] AM-IV sets can be created by performing the similar AM transformation as described in footnote 19. Hence, we can construct up to $T(k_1 + \ell)$ additional instruments, where $\ell = 5$ in CEEP and $\ell = 6$ in PC. Again, due to low cross-variations of $(\mathbf{QX}_1)^*$ and $\left(\mathbf{Q}\hat{\boldsymbol{\xi}}\right)^*$, we only consider subsets of $T(k_1 + \ell)$ to avoid collinearity. Beginning with the first T years we include as many instruments as possible. The final selection is made on the basis of the Sargan test results. Further, we do not consider the alternative set of instruments $(\mathbf{QX}_2)^*$ proposed by BMS because the dummies CEE and EMU in $\mathbf{X}_2 = (TGDP, SIM, CEE, EMU)$ do not vary across country-pairs over a number of years (EMU is 0 before than 1999 and CEE is always 1 after 1995), leading to perfect multicollinearity.

which is a plausible finding given that the impact of RLF on total trade flows (the sum of inter- and intra-industry trades) might not be unambiguous.[22] Secondly, the effect of similarity turns out to be higher. Combined together, these results clearly confirm that the intra-industry trade has become the main part of the total EU trade.[23] Thirdly and importantly, the impacts of EMU and CEE are still significant, but they become substantially smaller. The Euro impact drops sharply from 0.141 to 0.036 and 0.045 for CCEP and PC estimators. On the other hand, the impact of EMU falls modestly from 0.31 to 0.23 for CCEP estimator, but substantially to 0.08 for PC estimator.

Turing to HT-IV and AM-IV estimation results for investigating the impacts of time-invariant regressors, we find that distance and language dummies have significantly negative and positive impacts whereas the border impact is still insignificant, a finding consistent with those in Table 2. Furthermore, the Hausman test does not reject the legitimacy of the AM-IV estimates as more efficient than HT-IV estimates. Notice that the CEEP and the PC estimation results are qualitatively similar except for two main differences: (i) the PC estimate of $TGDP$ is smaller and (ii) the PC estimated impact of RER is much stronger than the CEEP counterpart.

Table 4 about here

Table 5 reports the estimation results for the extended model with cross-section dependence and with heterogeneous linear trends. The current estimation results are qualitatively similar to those presented in Table 4. Importantly, however, we now observe that the CEEP and PC estimates are remarkably similar. First, the coefficient of $TGDP$ converges at 1.85-1.9.[24] Secondly, the Euro impacts are significantly estimated at 0.033 and 0.03, substantially smaller than 0.14 reported in Table 2, and this magnitude is consistent with broad evidence compiled by Baldwin (2006) and more recent studies as reviewed in Section 2. Thirdly, the impacts of CEE stay at around 0.1, significantly lower than the figures obtained from the model without accommodating cross-section dependence. Finally, focussing on more efficient AM-IV estimates as confirmed by the Hausman test

[22] This is because the larger difference may result in the higher (lower) volume of inter- (intra-) industry trade.

[23] The share of the intra-trade has increased from 37.2% in 1960 to around 60% from 1990 onwards (see Table 1).

[24] Serlenga (2005) estimates coefficients on GDP_h and GDP_f, separately, using the triple index model, where h and f indicate home and foreign countries, and finds that the sum of coefficients on GDP_h and GDP_f are close to the coefficient on TGDP_{hf} obtained from the current double index model. GDP_f is a measure of the extent that exports are absorbed as the foreign economy grows whilst GDP_h is a measure of the size of the (domestic) economy. In this regard, theory predicts that each coefficient is close to a unit elasticity.

results, we find that distance and common language dummies exert significantly negative and positive impacts on trade. Again, the border impact is insignificant. In light of our *a priori* expectations and survey evidence reviewed in Section 2, we therefore conclude that the estimation results obtained by explicitly taking into account cross-section dependence and heterogeneous time trends, are mostly sensible, though the improvements can be made mostly by the former.

Table 5 about here

As a further robustness check we will investigate the two more important issues - the time-varying trade effects of bilateral resistance terms including the border effects and the issue of trade diversion - on the basis of our preferred estimation results summarised in Table 5.

Surprisingly, most existing studies neglect an important issue of assessing the effect of currency union on trade through bilateral resistance channels. In this regard we propose to test the validity of the following hypothesis: if the Euro had a positive effect on internal European trade (by reducing overall trade costs), this might have caused a decrease in trade impacts of bilateral trade barriers, especially the border effects (e.g. Cafiso, 2010). This will provide an alternative way to testing the Euro effect on trade integration. Consequently, we will check whether the trend line of coefficients of bilateral resistance proxies are more downward-sloping after 1999 than before 1999, in which case we deduce a positive effect of the Euro in terms of European Integration. To address this issue we re-estimate the model, (7), by the cross-section regressions for each time period such that we can estimate the time-varying coefficients of $\boldsymbol{\gamma}$. Notice that this estimation can be easily conducted within our framework after consistently estimating $\hat{d}_{it}$ in (7) by either CEEP or PC estimation. Here, we consider two options - (potentially biased) OLS and AM-IV. In the latter case we employ k_1 instruments at each time period (see footnote 19).

Figure 1 displays the time-varying estimation results obtained by OLS and AM-IV. It is clear in both cases that the downward slopes of both border and language effects are steeper after than 1999 than before 1999. It is also remarkable to observe that their decreases turn sharp and are monotonic from the year when the Euro was launched. First, the border impacts display a rather fluctuating but slightly declining pattern before 1999, after which they monotonically decline, suggesting that the Euro effect may help to reduce border-linked trade costs.[25] Next, the language impact had been overshooting until the mid 1970s, and then followed a

[25] Using the shorter sample period (1995-2003), Cafiso (2010) documents an opposite finding that the border effect decreased by 25.6% during 1995-99 and by 10.6% during 1999-2003, and concludes that "there was no acceleration in the reduction of border-linked costs after the Euro." Notice however that his approach is based on an alternative strategy by McCallum (1995), and derives the border effects by appending the NT national trade observations to to bilateral

general declining trend until recently.[26] In particular, the language impact starts to decline around 1990 and its downward slope becomes steeper just after the Euro introduction in 1999. This downward trend may reflect the progressive lessening of restrictions on labor mobility within EU that encouraged migration and thus reduced the relative importance of cultural and linguistic trade barriers.[27] In fact, net migration (immigrants minus emigrants) in the EU registers an increasing trend after 1990, probably capturing the effect of the Maastricht Treaty in 1993 (World Bank, 2012).[28] Finally, turning to the distance effects on trade, we find that its impacts have been more or less stable over the full sample period, showing no clear pattern of downward trending. This is generally consistent with findings in the meta-study of a large number of estimated distance effects conducted by Disdier and Head (2008), who document that the trade elasticity with respect to distance does not decline over time, but rather increases. This may confirm that the notion of the death of distance has been difficult to identify in present-day trade data (Jacks, 2009). Overall, we might conclude that the introduction of the Euro helps to reduce trade effects of bilateral trade barriers and promote more integration among the EU countries by eliminating exchange rate-related uncertainties and transaction costs.

Figure 1 about here

Next, in order to examine the (indirect) trade diversion effects, we follow Micco et al. (2003) and Flam and Nordström (2006), and construct the dummy variable, TD, which is equal to 1 when only one country in the pair is the member of the Euro. Adding this dummy in the final specification used in Table 5, we find that the estimation results (unreported to save the space) are qualitatively similar to those presented in Table 5. The (direct) Euro impacts are still significant at 0.03-0.04 while the impacts of CEE stay at around 0.1. As expected,[29] we find that

trade flows. Hence, it would be interesting to investigate the relationship between the Euro introduction and the border effect (interpreted as integration effect in terms of trade costs) using the framework advanced by McCallum (1995). However, as he employs the panel regression with conventional two-way fixed effects without accommodating cross-section dependence, we predict that his estimation results may be likely to be misleading, see also footnote 17.

[26] Egger and Lassmann (2012) provides a meta-analysis based on 701 language effects collected from 81 academic articles. On average, a common language increases trade flows by 44%.

[27] Immigrants promote trade with their country of origin, e.g. Rauch and Trindade (2002).

[28] The Treaty of Maastricht in 1993 introduced the concept of citizenship of the European Union which confers every Union citizen a fundamental and personal right to move and reside freely without reference to an economic activity.

[29] Depending on trade effects mechanism, the sign of this dummy coefficient is expected to be mostly ambiguous. If adoption of the Euro operated in the same way as a preferential trade liberalisation (i.e. leading to the supply switching), the TD coefficient should be negative. Alternatively, an individual country's adoption of the Euro may make it a more open economy, which will boost its trade with all other nations.

the (indirect) trade diversion effect is slightly positive but statistically significant.[30] Our evidence is generally in line with the results presented in Micco et al. (2003).

5 Conclusions

The investigation of unobserved multilateral resistance terms in conjunction with omitted trade determinants has recently assumed a prominent role in the literature on the Euro's trade effects (Baldwin, 2006). In this paper we follow recent developments in panel data studies (Ahn et al., 2001, Pesaran, 2006; Bai, 2009), and extend the cross-sectionally dependent panel gravity models advanced by Serlenga and Shin (2007). The desirable feature of this approach is to control for time-varying multilateral resistance, trade costs and globalisation trends explicitly through the use of both observed and unobserved factors, which are modelled as (strong) cross-sectionally correlated. Furthermore, this approach allows us to consistently estimate the impacts of (potentially endogenous) bilateral trade barriers such as the border and the common language dummies through combining the CEEP and the PC estimators with HT, AM and BMS IV estimators.

Applying the proposed cross-sectionally dependent panel gravity model to the dataset over the period 1960-2008 (49 years) for 91 country-pairs amongst 14 EU member countries, we obtain stylised findings as follows: Firstly, as expected, the sum of home and foreign country GDPs significantly boosts trade while a depreciation of the home currency increases trade flows. Secondly, the impact of difference in relative factor endowments is no longer significant whilst the effect of similarity turns out to be substantially larger. This suggests that similarity (in terms of countries' GDP rather than relative factor endowments) helps to ease the integration process by capturing trade ties across countries. Thirdly, the impacts of both distance and common language on trade are significantly negative and positive, attesting the validity of these proxies to capture bilateral trade barriers, though the border impact is no longer significant. Finally and importantly, the Euro's trade effect amounts to 3-4% only, even after controlling for trade diversion effects. We also find that the custom union effect is substantially reduced to 10% from 31% (without accommodating cross-section dependence). These small effects of both currency and custom unions provide a support for the thesis that the trade increase within the Euro area may reflect a continuation of a long-run historical trend, probably linked to the broader set of EU's economic integration policies and institutional changes, e.g. Berger and Nitsch (2008), and Lee (2012). While the advent of the Euro might be a necessary condition for the European integration process to continue beyond the single market agenda in the early 1990s,

[30] The CEEP and PC estimates are 0.03 (0.017) and 0.002 (0.017), where figures inside bracket stand for the standard errors.

the Euro's repercussions on trade are difficult to understand without taking proper account of the process of the underlying European institutions. An obvious policy implication is that countries considering joining the Euro would benefit from the ongoing process of integration, but should also be wary of regarding promises of an imminent acceleration of intra-area trade.

It is worth mentioning some of the avenues for further research opened by this paper. Firstly, we aim at analysing the relevant nexus between the Euro and trade imbalances among Euro members. Berger and Nitsch (2010) find that trade imbalances have been widening considerably after the introduction of the Euro. An application of our generalised approach will be expected to shed further lights on this politically sensitive issue. Secondly, it would be worthwhile to investigate in details the relationship between our factor-based approach and the spatial-based techniques developed by Behrens et al. (2012) for capturing unobserved multilateral resistance and trade costs by controlling for cross-sectional interdependence. We will examine this issue further by following and modifying the recent work of Mastromarco et al. (2013).

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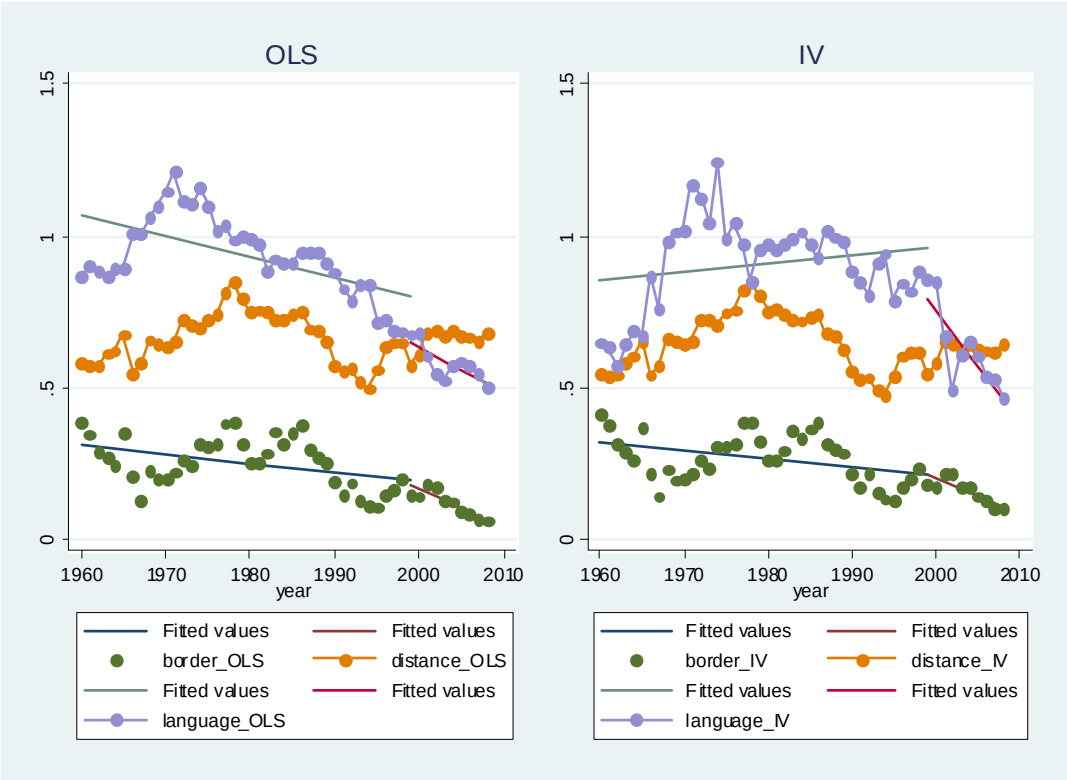
Figure 1: Time-varying estimation of the trade effects of bilateral trade barriers



Notes: This figure shows the time-varying estimation of the trade effects of bilateral trade barriers with the relative fitted values. The left panel plots OLS estimates, and the right panel displays AM-IV estimates using the AM set of instruments (see footnotes 19 and 21). The time-varying coefficients are obtained in two steps: first, we estimate model (1)-(2) by CCEP including heterogeneous trends as in Table 5, and then estimate (7) by cross-section regressions for each time period.

Table 1: EU trade shares and growths

Panel A	1960[1]	1970[2]	1980[3]	1990[4]	2000[5]	2008[6]
Share of US on Extra-EU trade	16.5	26.3	33.8	19	21.9	15.1
Share of Intra-EU on EU trade	37.2	49.8	50.5	59.7	61.7	61
Share of Export on Intra-EU trade	52.4	51.6	51.1	49.7	51.2	50.1
Panel B	60/70	70/80	80/90	90/00	00/08	
Average Growth of GDP	8.9	16.4	7.8	3.5	2.5	
Average Growth of Intra-EU trade	11.5	17.3	9.3	5.8	5.6	
Average Growth of Total EU trade	10.3	20.1	7.2	3.9	8.1	
Average Growth of Bilateral Exchange Rate	0.12	7.9	-1.4	-3.7	-2.3	

Notes: 1 refers to EU6 (Belgium, France, Germany, Italy, Luxemburg, Netherlands) from 1960 to 1969; 2 refers to EU6 from 1970 to 1973 and EU9 (EU6 plus Denmark, Ireland and UK) from 1973 to 1979; 3 refers to EU9 in 1980, EU10 (EU9 plus Greece) from 1981 to 1985, and EU12 (EU10 plus Portugal and Spain) from 1986 to 1989; 4 refers to EU12 from 1990 to 1994 and EU15 (EU12 plus Austria, Finland and Sweden) from 1995 to 1999; 5 refers to EU15 from 2000 to 2001; 6 refers to EU15 from 2001 to 2004 and EU25 (EU15 plus Cyprus, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Slovakia and Slovenia) and EU27 (EU25 plus Romania and Bulgaria) from 2007 to 2008, respectively. Sources: Statistical Yearbook, Eurostat (1997) and Trade Policy Review of the European Union: A Report by the Secretariat of the WTO, WTO (2002), Unctatd (2012), World Bank (2012).

Table 2: Estimation results without cross section dependence

	POLS	FE	RE	HT	AM
gdp	1.7002**	2.0271**	2.0123**	2.0269**	2.0195**
	[0.01]	[0.02]	[0.02]	[0.02]	[0.02]
sim	0.8234**	1.2044**	1.1820**	1.2063**	1.1830**
	[0.02]	[0.06]	[0.05]	[0.06]	[0.05]
rfl	0.0616**	0.0464**	0.0478**	0.0458**	0.0471**
	[0.01]	[0.01]	[0.01]	[0.01]	[0.01]
rer	-0.0634**	0.0273**	0.0221*	0.0252**	0.0239**
	[0.01]	[0.01]	[0.01]	[0.01]	[0.01]
euro	0.2682**	0.1413**	0.1472**	0.1412**	0.1449**
	[0.03]	[0.02]	[0.02]	[0.02]	[0.02]
cee	0.3568**	0.3055**	0.3114**	0.3048**	0.3087**
	[0.02]	[0.02]	[0.02]	[0.02]	[0.02]
		OLS			
distance	-0.7812**	-0.605**	-0.6128**	-0.5219**	-0.6071**
	[0.02]	[0.02]	[0.10]	[0.15]	[0.12]
border	0.1366**	0.047	0.0502	-0.286	0.0383
	[0.04]	[0.03]	[0.21]	[0.43]	[0.25]
language	0.5091**	0.855 **	0.8373**	1.8064+	0.8702**
	[0.04]	[0.04]	[0.25]	[1.04]	[0.31]
Sargan				$\chi^2_1 = 1.91$	$\chi^2_{61} = 66.21$
p-value				0.166	0.302
Hausman		H:$\chi^2_6 = 32.5$			H1:$\chi^2_9 = 13.85$
p-value		0.000			0.127

Notes: Using the annual data over 1960-2008 for 91 country-pairs, we estimate the model (1)-(2) without cross section dependence. The dependent variable is the logarithm of real total trade flows and the regressors are $\mathbf{x}'_{it} = \{RER, TGDP, RLF, SIM, CEE, EMU\}_{it}$ and $\mathbf{z}_i = \{DIS, BOR, LAN\}_i$. POLS stands for the pooled OLS estimator, FE for fixed effects estimator and RE for random effects estimator, respectively. Figures in $[\cdot]$ indicate the standard error. **,* and + denote 1, 5, and 10 per cent level of significance, respectively. The cross dependence Pesaran (2004) test for cross-correlation rejects the null no cross dependence at 1 per cent level of significance. H denotes the Hausman statistic under the null of no correlation between explanatory variables and individual effects. H1 denotes the Hausman statistic testing for the legitimacy of the AM estimates above the corresponding p-values. Instrumental variables in the HT and AM estimates are IV $= \{RER_{it}, RLF_{it}\}$. Sargan denotes the statistic testing for the validity of over-identifying restrictions above the corresponding p-values. .

Table 3: Estimation results without cross section dependence and with heterogeneous trends

	POLS	RE	FE	HT	AM
constant	0.0006	0.005	-2.922 **	-3.606 *	-3.005 **
	[0.007]	[0.019]	[0.969]	[1.352]	[1.068]
gdp	1.788 *	1.785 *	2.031 **		
	[0.01]	[0.012]	[0.018]		
sim	0.862 *	0.86 *	1.177 **		
	[0.023]	[0.026]	[0.057]		
rfl	0.067 *	0.061 *	0.041 **		
	[0.008]	[0.008]	[0.007]		
rer	-0.028 *	-0.027 *	0.022 *		
	[0.007]	[0.008]	[0.01]		
euro	0.093 *	0.068 *	0.114 **		
	[0.028]	[0.028]	[0.017]		
cee	0.234 *	0.242 *	0.311 **		
	[0.021]	[0.022]	[0.015]		
			OLS		
distance	-0.669 *	-0.66 *	-0.601 **	-0.508 *	-0.593 **
	[0.018]	[0.02]	[0.134]	[0.183]	[0.144]
border	0.044	0.043	0.041	-0.329	0.044
	[0.052]	[0.058]	[0.272]	[0.415]	[0.207]
language	0.88 *	0.891 *	0.861 *	1.921 +	0.932 +
	[0.063]	[0.071]	[0.328]	[1.196]	[0.352]
trend	0.009	0.008	0.18 *		
	[0.013]	[0.014]	[0.013]		
Sargan				$\chi^2_1 = 1.815$	$\chi^2_{58} = 46.84$
p-value				0.178	0.853
Hausman			H:$\chi^2_6 = 36.1$		H1: $\chi^2_3 = 5.54$
p-value			0		0.211

Notes: We estimate the model (1)-(2) without cross section dependence and with heterogeneous trends. Trend shows the Mean Group estimates the heterogeneous trend coefficients. Instrumental variables in the HT and AM estimates are IV = $\{RER_{it}, RLF_{it}\}$. See also notes to Table 2.

Table 4: Estimation results with cross section dependence

	CCEP	HT	HT1	AM	AM1	PC	HT	HT1	AM	AM1
gdp	2.06 **					1.487 **				
	[0.047]					[0.052]				
sim	1.512 **					1.458 **				
	[0.098]					[0.072]				
rfl	0.005					-0.001				
	[0.014]					[0.004]				
rer	0.031 **					0.142 **				
	[0.008]					[0.017]				
euro	0.036 **					0.045 **				
	[0.007]					[0.015]				
cee	0.231 **					0.08 **				
	[0.007]					[0.012]				
	OLS					OLS				
constant	-2.745 **	-3.566 **	-3.822 **	-2.87 **	-2.709 **	5.921 **	4.181 **	3.498 **	5.801 **	5.694 **
	[0.919]	[1.403]	[1.249]	[1.137]	[1.118]	[1.233]	[1.284]	[1.401]	[0.917]	[0.907]
distance	-0.575 **	-0.464 *	-0.429 **	-0.56 **	-0.578 **	-0.851 **	-0.615 **	-0.522 **	-0.837 **	-0.825 **
	[0.127]	[0.189]	[0.168]	[0.154]	[0.15]	[0.17]	[0.176]	[0.193]	[0.128]	[0.126]
border	0.056	-0.389	-0.527	0.019	0.026	0.31	-0.632	-1.002	0.302	0.275
	[0.258]	[0.475]	[0.43]	[0.26]	[0.25]	[0.346]	[0.722]	[0.804]	[0.231]	[0.234]
language	0.969 **	2.243 *	2.64 **	1.109 **	1.017 **	0.622	3.321 +	4.379 *	0.721 +	0.842 *
	[0.311]	[1.271]	[0.992]	[0.414]	[0.388]	[0.417]	[1.724]	[1.683]	[0.415]	[0.438]
$RERT_t$	-0.006 *					-0.002				
	[0.002]					[0.001]				
Sargan		$\chi^2_1 = 0.67$	$\chi^2_6 = 10.07$	$\chi^2_{47} = 53.23$	$\chi^2_{67} = 79.02$		$\chi^2_1 = 15.64$	$\chi^2_7 = 17.65$	$\chi^2_{53} = 58.9$	$\chi^2_{58} = 68.21$
p-value		0.411	0.121	0.246	0.149		0	0.014	0.268	0.169
Hausman				H1: $\chi^2_3 = 2.38$	H1:$\chi^2_3 = 5.68$				$\chi^2_3 = 4.007$	$\chi^2_3 = 7.145$
p-value				0.497	0.128				0.405	0.128

Notes: We estimate the model (1)-(2) with cross section dependence, CCEP denotes the Pesaran (2006) CCEP estimation whereas PC denotes the PC estimator proposed by Bai (2009). In the CCEP $\mathbf{f}_t = \left\{RERT_t, \overline{TGDP}_t, \overline{SIM}_t, \overline{RLF}_t, \overline{CEE}_t\right\}$ whereas in PC $\mathbf{f}_t$ are six factors extracted using Bai and Ng (2002) procedure plus $\{RERT_t\}$. $RERT_t$ shows the Mean Group estimates of the heterogeneous $RERT_t$ coefficients. For the HT and the AM estimates we consider the following sets of instruments: IV = $\{RER_{it}, RLF_{it}\}$ and IV1 = $\left\{IV, \hat{\boldsymbol{\lambda}}_i \mathbf{f}_t\right\}$. See also notes to Table 2.

Table 5: Estimation results with cross section dependence and with heterogeneous trends

	CCEP	HT	HT1	AM	AM1	PC	HT	HT1	AM	AM1
gdp	1.85 **					1.885 **				
	[0.027]					[0.052]				
sim	1.531 **					1.298 **				
	[0.029]					[0.083]				
rfl	-0.003					-0.003				
	[0.003]					[0.004]				
rer	0.058 **					0.062 **				
	[0.008]					[0.016]				
euro	0.033 **					0.03 *				
	[0.003]					[0.014]				
cee	0.099 **					0.106 **				
	[0.003]					[0.012]				
	OLS					OLS				
constant	0.994	-0.275	-0.073	1.021	0.96	4.723 **	3.082 **	2.909 *	4.538 **	4.534 **
	[0.908]	[1.25]	[1.145]	[1.051]	[1.033]	[0.991]	[1.373]	[1.245]	[0.999]	[0.974]
distance	-0.69 **	-0.593 **	-0.546 **	-0.692 **	-0.684 **	-0.679 **	-0.457 *	-0.434 *	-0.66 **	-0.653 **
	[0.125]	[0.17]	[0.155]	[0.143]	[0.141]	[0.136]	[0.187]	[0.169]	[0.135]	[0.132]
border	0.154	-0.235	-0.424	0.133	0.106	0.116	-0.772	-0.866	0.104	0.009
	[0.254]	[0.44]	[0.428]	[0.241]	[0.242]	[0.278]	[0.697]	[0.643]	[0.209]	[0.215]
language	0.836 **	1.951 +	2.491 *	0.849 *	0.935 *	0.835 *	3.379 *	3.646 *	0.992 *	1.134 **
	[0.307]	[1.177]	[0.988]	[0.403]	[0.391]	[0.335]	[1.648]	[1.358]	[0.365]	[0.366]
$RERT_t$	-0.006 *					-0.001				
	[0.002]					[0.001]				
Sargan		$\chi^2_1 = 1.41$	$\chi^2_7 = 11.21$	$\chi^2_{47} = 56.03$	$\chi^2_{46} = 57.35$		$\chi^2_1 = 5.41$	$\chi^2_9 = 11.19$	$\chi^2_{48} = 54.71$	$\chi^2_{58} = 63.85$
p-value		0.234	0.13	0.172	0.122		0.019	0.262	0.235	0.278
Hausman				H1:$\chi^2_3 = 0.01$	H1:$\chi^2_3 = 3.04$				H1:$\chi^2_3 = 3.01$	H1:$\chi^2_3 = 5.88$
p-value				0.999	0.55				0.509	0.21

Notes: We estimate the model (1)-(2) with cross section dependence and with heterogneous trends. For the HT and AM estimations we consider the following sets of exogenous variables: IV = $\{RER_{it}, RLF_{it}\}$ and IV1 = $\left\{IV, \hat{\boldsymbol{\lambda}}_i \mathbf{f}_t\right\}$. See also notes to Tables 2 and 4 .