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A PANEL COINTEGRATION STUDY OF THE EURO EFFECT ON TRADE

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Abstract

Bun and Klaassen (2007) investigate the impact of the introduction of Euro on bilateral trade. Accounting for deterministic trends in the residuals of the gravity equation they estimate an Euro effect of about 3%, smaller than previous estimates in the range of 5% to 40%. In this paper we revisit their data using methods recently advanced in the analysis of non-stationary panel data with cross-sectional dependence. Using several panel unit root tests we find strong evidence that (the log of) bilateral trade, as well as the product of GDP and GDP per capita have unit roots. However, we find cointegration between these variables using the cointegration test of Gengenbach et al. (2006) and the error correction tests proposed by Gengenbach et al. (2008). Employing the common correlated effects (CCEP) estimator of Pesaran (2006) and the continuously updated (CUP) estimator of Bai et al. (2009), we obtain estimates of the cointegrating vector and estimates of the Euro effect on bilateral trade. Our estimates vary between models and estimators but seem to support the findings of Bun and Klaassen (2007).

Keywords: Gravity models, trade, panel cointegration, common factors.

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1 Introduction

The gravity model of trade has been widely used in the empirical literature to study the effect of various policy decisions (see e.g. Tinbergen, 1962, for an early application). In its original form the gravity model explains trade flows between country pairs as being proportional to their national income and inverse proportional to their distance. It has been successful in empirical studies of the effect of various variables on trade and also received several theoretical foundations (see e.g. Anderson, 1979; Bergstrand, 1985, 1989). The original gravity model has been augmented by numerous additional variables. Population size or GDP per capita have been added as additional measures of mass. Features of the geographic location such as longitude/latitude or dummy variables for landlocked or island nations. Furthermore, variables measuring whether country pairs share certain aspects have been added, such as common border, common language or a common colonial history.

One particular area of interest is the impact of currency unions on trade. Rose (2000) first estimates the effect of currency unions on trade in a cross-sectional study and finds that adopting a currency union leads to a 200% increase in bilateral trade. Glick and Rose (2002) and Frankel and Rose (2002) perform panel data studies of the common currency effect on trade and obtain similarly large estimates. It is also found that the increase in trade between countries sharing a common currency is not due to a deviation of trade from other partners but due to an increase in total trade. Frankel and Rose (2002) also investigate the effect of trade on income and find that a 1% increase in total trade leads to an increase in income per capita by about one-third of a percent. These very high estimates have led to a controversy in the empirical literature. In particular, most observations on currency unions in Rose's data comes from poorer or small countries or dependencies. This has led to the question whether the result applies to bigger countries such as the members of the European Monetary Union (EMU) (see the discussion in e.g. Glick and Rose, 2002; Frankel and Rose, 2002).

In 1999, 11 countries of the EMU (Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, the Netherlands, Portugal and Spain) introduced the Euro as a common currency. Greece joined the Euro in 2001, after the initial launch but before the introduction of Euro notes and coins. Since then, also Slovenia, Cyprus, Malta and Slovakia have joined the Euro. At the same time, other member states of the European Union (EU) have decided not to adopt the Euro, namely Denmark, Sweden and the United Kingdom. It is therefore an important question if there is a Euro effect on bilateral trade, i.e if there is an economic reward for joining the common currency. Furthermore, the introduction of the Euro itself was preceded by the European Monetary System (EMS) aimed at aligning the European exchange rates. The EMS might have had a mitigating effect on the introduction of the Euro on bilateral trade. But it might be interesting to investigate whether there is an additional benefit of a common currency over (relative) exchange rate stability. Early studies report an

Euro effect between 5% and 40%. Micco et al. (2003) estimate an increase in bilateral trade due to the introduction of the Euro between 8% and 16% when compared to trade between non-EMU countries. Flam and Nordström (2003) estimate the Euro effect between 5% and 15%. Bun and Klaassen (2002) estimate a dynamic panel data model of the gravity equation and find a long-run Euro effect of about 38%, where the immediate effect is estimated at about 3.9% increase in 1999. Rose and Stanley (2005) perform a meta analysis of the results of 34 studies of the effect of currency unions on trade. Although they find evidence of publication bias, they also find evidence that currency unions have a significant positive effect on bilateral trade, and obtain a combined estimate of the trade effect between 30% and 90%.

Bun and Klaassen (2007) estimate the gravity equation allowing for country pair specific time trends to account trending behavior observed in the residuals. This reduces the estimate of the Euro effect on bilateral trade to about 3%. They also investigate whether the data is non-stationary and find unit roots in the (log of) bilateral trade, GDP and GDP per capita using the panel unit root test of Harris and Tzervalis (1999) and the panel stationarity test of Hadri (2000). Furthermore, they find cointegration between these variables using the panel cointegration test of Pedroni (1999) and estimate the coefficients using the dynamic OLS (DOLS) estimator of Mark and Sul (2003). However, the employed methods assume that the data is cross-sectionally independent, an assumption unlikely to hold bilateral trade data.

Cross-sectional dependence has received much attention recently in the literature on non-stationary panel data as the assumption of cross-sectional independence is unlikely to hold in many data sets. It is of particular interest here as bilateral trade data by construction is highly cross-sectionally related. Furthermore, the gravity model itself implies spatial dependence in the data due to the hypothesized effect of distance on trade. Several new panel unit root or cointegration tests have been proposed that allow for cross-sectional dependence in the form of common factors. See for example Breitung and Pesaran (2008) for an overview of the literature and Gengenbach et al. (2009) for a comparison of panel unit root tests. We use the panel unit root tests proposed by Pesaran (2007), Moon and Perron (2004), Breitung and Das (2008), Sul (2007), Bai and Ng (2004) and Palm et al. (2008) to test whether the variables entering the gravity model are non-stationary. We then test for cointegration between the variables using the panel no-cointegration test proposed by Gengenbach et al. (2006) and the panel no error-correction tests suggested by Gengenbach et al. (2008) to investigate whether the variables are cointegrated. Both approaches allow for persistent cross-sectional dependence in the data in form of unobserved common factors. We estimate the cointegrating vector using the CUP estimator of Bai et al. (2009) and the CCEP estimator of Pesaran (2006) and obtain an estimate of the Euro effect on bilateral trade.

The remainder of this paper is organized as follows. Section 2 describes the data, summarizes the main findings of Bun and Klaassen (2007) and presents a brief overview of other

studies of the Euro effect on trade. Section 3 presents the results of the panel unit root and panel cointegration tests. We obtain estimates of parameters of the cointegrated gravity model in Section 4. Section 5 concludes.

2 Data and previous studies

We use the data set of Bun and Klaassen (2007) which contains data on all bilateral combinations in a panel of 19 countries, namely the 15 member countries of the EU prior to the 2004 expansion as well as Norway, Switzerland, Canada, Japan and the US. The data for Belgium and Luxembourg is combined because trade data for these countries is only available for the Belgium-Luxembourg Economic Union. The data spans the time period between 1967 and 2002. Thus we have a balance panel with $N=171$ country pairs and $T=36$ time series observations.

The data set includes the following variables. $TRADE_{ijt}$ is the log of real bilateral trade between countries i and j at time t , where real bilateral trade is measured as the sum of nominal bilateral exports and imports in US dollars divide by the US producer price index. GDP_{ijt} is the log of the product of countries' real GDP. $GDPCAP_{ijt}$ measures the log of the product of the countries' real GDP per capita. Furthermore, 2 dummy variables are included in the data, namely $EURO_{ijt}$ which is 1 if both countries have adopted the Euro at time t and FTA_{ijt} which is 1 if both countries have a free trade agreement at time t . The model estimated by Bun and Klaassen (2007) is given by the following equation.

$$TRADE_{ijt} = \beta_1 GDP_{ijt} + \beta_2 GDPCAP_{ijt} + \delta_1 EURO_{ijt} + \delta_2 FTA_{ijt} + \eta_{ij} + \tau_{ij} \cdot t + \lambda_t + \varepsilon_{ijt}, \quad (1)$$

where η_{ij} is a country pair specific fixed effect, λ_t is a common time effect, $t_{ij} \cdot t$ is a country pair specific time trend and ε_{ijt} is the error term. The current model does not include distance between countries as a dependent variable. Nevertheless, country pair specific fixed effects will account for part of the distance effect, and any time invariant measure of distance would be removed by the within transformation. Furthermore, Pesaran and Tosetti (2009) show that cross-sectional averages are well suited to account for spatial dependence.

Table 1 presents the estimates of the parameters of Equation (1) obtained by Bun and Klaassen (2007). When not allowing for time trends the estimated Euro effects are 51% and 45% for the LSDV and DOLS estimators, respectively. These estimates are above effects reported in earlier studies. However, the effect is reduced to only about 3% for both estimators when time trends are included in the model. The reduced estimate is robust to various other specifications employed by Bun and Klaassen (2007).

Similar to the Euro effect, the estimated effect of a free trade agreement between countries is also reduced when accounting for time trends in the estimation. The coefficient of GDP_{ijt}

Table 1: Estimation results Bun and Klaassen (2007)

	LSDV		DOLS	
	no trends	trends	no trends	trends
δ_1	0.410	0.032	0.374	0.034
δ_2	0.41	0.06	0.38	0.05
β_1	1.41	0.70	0.59	0.94
β_2	-0.68	-0.23	0.20	-0.49

Notes: “No trends” indicates that τ_{ij} is set to 0 in (1). LSDV gives results for LSDV-type estimates from Bun and Klaassen (2007, Table 2, p. 480). DOLS give the estimates from Bun and Klaassen (2007, Table 5, p. 491).

is positive, but there are some differences in the size of the estimated effect, ranging from 0.59 to 1.41. The coefficient of $GDPCAP_{ijt}$ is negative for 3 out of the 4 case.

Bun and Klaassen (2002) estimate a dynamic gravity model for real exports. They have annual data ranging from 1965 to 2001 for the 15 EU member countries before the 2004 expansion as well as on Canada, Japan and the US. They estimate an ADL model using the LSDV estimator and obtain a Euro effect estimate of about 4% in the first year with an estimated long-run effect of about 40%. However, they do not consider models with deterministic or stochastic trends. Flam and Nordström (2003) estimate the Euro effect in a panel of 20 OECD countries with annual observations from 1989 to 2002. The panel includes data on the 10 countries (combining the data on Belgium and Luxembourg) that entered the EMU in 1999 as well as Australia, Canada, Denmark, Japan, New Zealand, Norway, Sweden, Switzerland, the UK and the US. They detect a break associated with the introduction of the Euro already in 1998. For the period of 1998 to 2002, they estimate an average increase in exports between EMU countries of about 15% compared to average exports in the 1989 to 1997 period. Flam and Nordström (2003) also detect a spill-over effect of the Euro. They estimate an increase in exports from the Euro zone to non-Euro countries of about 8% and an increase in exports to the Euro zone from non-Euro countries of approximately 5%. Flam and Nordström (2003) also analyze sector specific data and find significant Euro concentrated to a few sectors. Micco et al. (2003) consider two data sets in their analysis: a panel of 22 industrialized countries included in the IMF’s Directions of Trade Statistics data set and a panel including only the 15 EU member countries prior to the 2004 expansion. Using different specification of the gravity equation, Micco et al. (2003) estimate a Euro effect between 4% and 26%. Furthermore, they also identify the Euro having an effect on trade starting in 1998, similar to Flam and Nordström (2003). They also estimate a spill-over effect of the Euro, increasing trade between Euro zone countries and non-Euro countries by up to 9%.

3 Panel unit root and cointegration tests

In this section we test for unit roots in the variables of the model and then test for cointegration between the non-stationary variables. We employ the tests proposed by Pesaran (2007), Moon and Perron (2004), Breitung and Das (2008), Sul (2007), Palm et al. (2008) and Bai and Ng (2004) to test for unit roots. We then use the no cointegration test advanced by Gengenbach et al. (2006) and the no error correction tests proposed by Gengenbach et al. (2008) to investigate whether the variables are cointegrated. We briefly outline the test procedures before presenting the results.

3.1 Panel Unit Root Tests

Considers a heterogenous, linear model for a balanced panel with N cross sectional units and T times series observations. In particular,

$$Y_{i,t} = (1 - \rho_i)\mu_i + \rho_i Y_{i,t-1} + u_{i,t}, \quad (2)$$

where the error term u_{it} has a common factor structure, such that

$$u_{i,t} = \gamma_i f_t + e_{i,t}.$$

Here, f_t is an unobserved common factor, γ_i is the corresponding factor loading and $e_{i,t}$ is an idiosyncratic error term independent across i and independent of the common factor. It is convenient to re-write (2) as

$$\Delta Y_{i,t} = \alpha_{0i} + \alpha_{1i} Y_{i,t-1} + \gamma_i f_t + e_{i,t}, \quad (3)$$

where $\alpha_{0i} = (1 - \rho_i)\mu_i$ and $\alpha_{1i} = (\rho_i - 1)$. Pesaran (2007) suggests to cross-sectionally augment the test equation (3) with cross-sectional averages of the first differences and the lagged levels to account for the cross-sectional dependence induced by a single common factor. The cross-sectionally augmented (CA)DF equation is then given by

$$\Delta Y_{i,t} = a_i + b_i Y_{i,t-1} + c_i \bar{Y}_{t-1} + d_i \Delta \bar{Y}_t + \varepsilon_{i,t}, \quad (4)$$

where $\bar{Y}_{t-1} = \sum_{i=1}^N Y_{i,t-1}$, $\Delta \bar{Y}_t = \sum_{i=1}^N \Delta Y_{i,t}$ and $\varepsilon_{i,t}$ is the regression error. The individual specific test statistic for the hypothesis $H_{0i} : \rho_i = 1$ for a given i is now the t-statistic of b_i in (4), denoted by $CADF_i$. The panel unit root for the hypothesis $H_0 : \rho_i = 1$ for all i against the heterogenous alternative $H_1 : \rho_i < 1$ for some i is given by the cross-sectional average of the $CADF_i$ tests, such that

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

For computational reasons, Pesaran (2007) advocates the use of a truncated version, \overline{CADF}^* , where for positive constants K_1 and K_2 such that $Pr[-K_1 < CADF_i < K_2]$ is sufficiently large values of $CADF_i$ smaller than $-K_1$ or larger than K_2 are replaced by the respective bound. Pesaran (2007) provides values for K_1 and K_2 as well as critical values for the test statistics obtained via stochastic simulation. In case the error terms or the common factor are serially the CADF equation (4) can be augmented by additional lags of \bar{Y}_{t-1} , $\Delta\bar{Y}_t$ and $\Delta Y_{i,t-1}$.

Moon and Perron (2004) propose two test statistics for the null hypothesis $H_0 : \rho_i = 1$ for all i against the heterogenous alternative $H_1 : \rho_i < 1$ for some i . They allow for k common factors in $u_{i,t}$. Their method relies on de-factoring the data by a projection onto the space orthogonal to that spanned by the factor loadings. They propose to estimate the factor loadings by method of principle components from the residuals of a pooled first stage regression,

$$\hat{u}_{i,t} = Y_{i,t} - \hat{\rho}_{pols} Y_{i,t-1},$$

where $\hat{\rho}_{pols}$ is the pooled OLS estimator of ρ_i in (2). The de-factored data is now given by

$$Y_{i,t}^* = Y_{i,t} - \hat{\gamma}_i \left(\sum_{j=1}^N \hat{\gamma}'_j \hat{\gamma}_j \right)^{-1} \sum_{j=1}^N \hat{\gamma}'_j Y_{j,t}.$$

The two test statistics suggested by Moon and Perron (2004) are based on a modified pooled estimator of ρ ,

$$\hat{\rho}^* = \left(\sum_{i=1}^N \sum_{t=2}^T (Y_{i,t-1}^*)^2 \right)^{-1} \left(\sum_{i=1}^N \sum_{t=2}^T Y_{i,t-1}^* Y_{i,t}^* - NT \hat{\phi}_e \right),$$

where $\hat{\phi}_e$ is the average estimated one-sided long-run covariance. The tests are given by

$$t_a^* = \frac{\sqrt{NT}(\hat{\rho}^* - 1)}{\sqrt{\frac{2\hat{\phi}_e^4}{\hat{\omega}_e^4}}} \quad (5)$$

and

$$t_b^* = \sqrt{NT}(\hat{\rho}^* - 1) \sqrt{\frac{1}{NT^2} \sum_{i=1}^N \sum_{t=2}^T (Y_{i,t-1}^*)^2 \left(\frac{\hat{\omega}_e}{\hat{\phi}_e^2} \right)}, \quad (6)$$

where $\hat{\omega}_e^2$ is the average estimated long-run covariance and $\hat{\phi}_e^4 = N^{-1} \sum_{i=1}^N \hat{\omega}_e^4$. Moon and Perron (2004) show that both test statistics have a standard normal limiting distribution.

Breitung and Das (2008) propose two tests for a unit root in (2), namely a generalized least squares (GLS) t-test, which is only feasible if $N < T$, and a robust t-test, t_{rob} . The later is given by

$$t_{rob} = \left(\sum_{t=2}^T Y'_{t-1} \hat{\Omega} Y_{t-1} \right)^{-\frac{1}{2}} \left(\sum_{t=2}^T Y_{t-1} \Delta Y_t \right), \quad (7)$$

where $Y_{t-1} = (Y_{1,t-1}, \dots, Y_{N,t-1})'$, $\Delta Y_t = (\Delta Y_{1,t}, \dots, \Delta Y_{N,t})'$ and $\hat{\Omega} = \sum_{t=2}^T \hat{u}_t \hat{u}_t'$, with $\hat{u}_t = (\hat{u}_{1,t}, \dots, \hat{u}_{N,t})'$ being the pooled OLS residuals. Breitung and Das (2008) show that t_{rob} converges to a Dickey-Fuller (DF) distribution under the null hypothesis $H_0 : \rho_i = 1$ for all i .

Palm et al. (2008) propose several bootstrap panel unit roots. They consider pooled Levin et al. (2002) type tests based on the pooled OLS estimate of ρ_i in (2) and group mean Im et al. (2003) type tests based on individual specific estimates of ρ_i . In particular, the pooled statistic is defined as

$$\tau_p = T(\hat{\rho}_{pols} - 1). \quad (8)$$

The group mean statistic is given by the following equation,

$$\tau_{gm} = N^{-1} \sum_{i=1}^N T(\hat{\rho}_i - 1), \quad (9)$$

where

$$\rho_i = \left(\sum_{t=2}^T Y_{i,t-1}^2 \right)^{-1} \left(\sum_{t=2}^T Y_{i,t-1} Y_{i,t} \right).$$

Palm et al. (2008) also consider τ_{med} which is given by T times the median of $(\hat{\rho}_i - 1)$, as the median might be more robust to outliers. Palm et al. (2008) propose a block bootstrap and show that it is asymptotically valid for a number of cross-sectional correlation models.

Bai and Ng (2004) consider a more general model than (2). In particular,

$$Y_{i,t} = \gamma_i F_t + E_{i,t}, \quad (10)$$

where F_t is a k -vector common factor and $E_{i,t}$ is the idiosyncratic component. They allow either F_t or $E_{i,t}$ to be non-stationary and propose to test them separately. As both common and idiosyncratic components are unobserved, Bai and Ng (2004) propose a consistent estimator. They apply the methods of principle components to the (demeaned) first differences of the data and re-accumulate the estimates to preserve the order of integration.

For the estimated idiosyncratic component, $\hat{E}_{i,t}$, Bai and Ng (2004) propose an ADF test to test for individual unit roots. To test the pooled unit root hypothesis that all $\hat{E}_{i,t}$ are non-stationary, Bai and Ng (2004) suggest a Fisher-type, using the correction proposed by Choi (2001) for the test of Maddala and Wu (1999). In particular, the test statistic is given by

$$P_{\hat{E}}^{c,\tau} = \frac{-2 \sum_{i=1}^N \log \pi_i - 2N}{\sqrt{4N}}, \quad (11)$$

where π_i is the p-value of the ADF test for the i -th cross section and c and τ denote the constant only or linear deterministic trend case, respectively. Bai and Ng (2004) show that $P_{\hat{E}}^{c,\tau}$ has a standard normal limiting distribution.

Bai and Ng (2004) propose several tests to select the number of independent stochastic trends, k_1 in the estimated common factors, \hat{F}_t . If a single common factor is estimated,

they propose an ADF test, $ADF_{\hat{F}}^{c,\tau}$. Bai and Ng (2004) show that the limiting distribution of $ADF_{\hat{F}}^{c,\tau}$ coincides with the Dickey-Fuller distribution for the respective cases. If more than one common factor is estimated, Bai and Ng (2004) propose an iterative procedure to select k_1 , similar to Johansen trace test for cointegration. Bai and Ng (2004) propose two modified Q statistics to test the hypothesis of $k_1 = m$ against the alternative $k_1 < m$ for m starting from \hat{k} . The procedure terminates if at any step $k_1 = m$ cannot be rejected. The two test statistics are denoted as $MQ_c^{c,\tau}$ and $MQ_f^{c,\tau}$, where the former uses a non-parametric correction to account for additional serial correlation while the later employs a parametric correction. Both statistics have a non standard limiting distribution and Bai and Ng (2004) provide critical values for several m .

Sul (2007) proposes recursive mean adjusted panel unit roots. He proposes a GLS test to test the hypothesis $H_0 : \rho_i = 1$ for all i against the heterogenous alternative $H_1 : \rho_i < 1$ for some i . However, the GLS test is not feasible if $T < N$. In case the data permits a Bai and Ng type representation as in (10), Sul (2007) proposes a recursive mean adjusted unit root test applied to the cross-sectional average of the data to test for a unit root in the common component. The test statistic is given by the FGLS t -test for $H_0 : \rho = 1$ in the following regression

$$\bar{Y}_t - \bar{C}_{t-1} = \rho(\bar{Y}_{t-1} - \bar{C}_{t-1}) + \sum_{j=1}^p \phi_j \Delta \bar{Y}_{t-j} + \varepsilon_{i,t}, \quad (12)$$

where $\bar{C}_{t-1} = \sum_{i=1}^N C_{i,t-1}$ with $C_{i,t-1} = (t-1)^{-1} \sum_{s=1}^{t-1} Y_{i,s}$. Sul (2007) provides simulated critical values for the test statistics, t_{crma} .

We apply the panel unit root tests described above to test for unit roots in $TRADE_{ijt}$, GDP_{ijt} and $GDPCAP_{ijt}$. The appropriate lag-lengths for tests is selected using the Akaike information criterion with a maximum $p = 4$. We use the Andrews and Monahan (1992) estimator employing the quadratic spectral kernel to estimate the nuisance parameters for the Moon and Perron (2004) tests. The number of common factors for the Moon and Perron (2004) and Bai and Ng (2004) test is estimated using the BIC_3 criterion of Bai and Ng (2002)¹ allowing for at most $k_{max} = 4$ factors. For the bootstrap tests of Palm et al. (2008)² we draw 10000 bootstrap samples. We use a fixed block length of $b = 6^3$. We allow for a linear trend in data.

The critical value for the \overline{CADF}^* test from is -2.56 at 5% level (see Pesaran, 2007, Table II(c)). With test statistic of -2.360 , -1.778 and -1.827 for $TRADE_{ijt}$, GDP_{ijt} and $GDPCAP_{ijt}$, respectively, we cannot reject the null hypothesis for all 3 panels. Using the asymptotic critical value of -1.645, the t_a^* test of Moon and Perron (2004) can reject the unit root null for $GDPCAP_{ijt}$ with a statistic of -9.620. The t_b^* test reject the null in all

¹The results are robust to using other selecting criterions and selecting different numbers of common factors.

²The author would like to thank Stephan Smeekes for providing the GAUSS codes for the test procedures.

³The results are robust for various block lengths $b = 1, \dots, \frac{T}{2}$.

three panels with values of -3.795, -11.402 and -11.982, respectively. The Breitung and Das (2008) t_{rob} test rejects the unit root null for $TRADE_{ijt}$ with a statistic of -4.606, using the asymptotic critical value of -3.41. Given the 5% asymptotic critical value of -1.86, the t_{crma} test of Sul (2007) rejects the unit root for all three panels. The $P_{\hat{E}}^\tau$ test of Bai and Ng (2004) cannot reject the unit root null using the asymptotic critical value of 1.645 for the estimated idiosyncratic component of either $TRADE_{ijt}$, GDP_{ijt} or $GDPCAP_{ijt}$. Estimating a single common factor for $TRADE_{ijt}$, the $ADF_{\hat{F}}^\tau$ test does not reject the unit root. Estimating 4 common factors in each panel for GDP_{ijt} or $GDPCAP_{ijt}$, both MQ_c^τ and MQ_f^τ cannot reject the null hypothesis that there are 4 independent stochastic trends. The critical values for the two statistics are -40.442 and -48.421, respectively (see Bai and Ng, 2004, Table I). The bootstrap panel unit root tests of Palm et al. (2008) cannot reject the unit root null in either of the three panels. For a block length of $b = 6$, the 5% bootstrap critical values for $TRADE_{ijt}$ are -12.491, -13.815 and -13.120 for τ_p , τ_{gm} and τ_{med} , respectively. For GDP_{ijt} , we obtain bootstrap critical values of -12.894, -13.627 and -13.475, while the critical values for the $GDPCAP_{ijt}$ panel are -12.777, -13.565 and -13.453.

As only the t_b^* of Moon and Perron (2004) and the t_{crma} of Sul (2007) are able to reject the unit root null for all three panels, there is strong evidence that the data is non-stationary⁴.

Table 2: Panel unit root and cointegration tests

Panel unit root tests								
variable	\overline{CADF}^*	τ_p	τ_{gm}	τ_{med}	t_a^*	t_b^*	t_{rob}	t_{crma}
$TRADE_{ijt}$	-2.360	-8.936	-9.438	-8.309	-1.225	-3.795 [†]	-4.606 [†]	-6.315 [†]
GDP_{ijt}	-1.778	-5.730	-5.975	-5.842	-1.539	-11.402 [†]	-2.154	-2.322 [†]
$GDPCAP_{ijt}$	-1.827	-5.841	-6.064	-5.899	-9.620 [†]	-11.982 [†]	-2.120	-2.361 [†]
variable	$P_{\hat{E}}^\tau$	$ADF_{\hat{F}}^\tau$	MQ_c^τ	MQ_f^τ				
$TRADE_{ijt}$	-10.49	-2.560	-	-				
GDP_{ijt}	-11.12	-	-17.07	-20.83				
$GDPCAP_{ijt}$	-11.07	-	-16.23	-21.67				
Panel cointegration tests								
	Z_ρ^p	Z_ρ^{gm}	Z_t^p	Z_t^{gm}	L_{trace}^F	$\bar{\tau}_{\alpha_i}^*$	$\bar{w}_{\delta_i}^*$	
	0.501	5.249	-4.949 [†]	-5.425 [†]	52.87 [†]	-5.051 [†]	62.60 [†]	

Notes: [†] denotes rejection at 5% level. Critical values are -2.56 for \overline{CADF}^* , -1.645 for t_a^* and t_b^* , -1.86 for -4.040 for t_{crma} , -3.41 for t_{rob} and $ADF_{\hat{F}}^\tau$, 1.645 for $P_{\hat{E}}^\tau$, -40.442 for MQ_c^τ and -48.421 for MQ_f^τ . Bootstrap 5% critical values for τ_p are -12.491, -12.894 and -12.777 for $TRADE_{ijt}$, GDP_{ijt} and $GDPCAP_{ijt}$, respectively. Bootstrap 5% critical values are -13.815, -13.627 and -13.565 for τ_{gm} and -13.120, -13.475 and -13.453 for τ_{med} .

The critical value for Z_ρ^p , Z_ρ^{gm} , Z_t^p and Z_t^{gm} is -1.645. The critical value for L_{trace}^F is 27.169. Critical values are -4.040 for $\bar{\tau}_{\alpha_i}^*$ and 28.203 for $\bar{w}_{\delta_i}^*$.

⁴Bayer and Hanck (2009) consider the possibility of combining several no cointegration tests. In principle, it should be possible to apply their method also to panel unit root tests.

3.2 Panel Cointegration Tests

Gengenbach et al. (2006) consider the problem of testing for no cointegration in a balanced panel with N cross-sections and T time series observations. For each cross-sectional unit, a $1 + m$ vector $Z_{i,t} = (Y_{i,t}, X'_{i,t})'$ is observed. Gengenbach et al. (2006) assume that both $Y_{i,t}$ and $X_{i,t}$ allow a Bai and Ng type representation as in (10). They propose to estimate the common and idiosyncratic components of the panels using the principle component estimator of Bai and Ng (2004). To test for no cointegration between the estimated idiosyncratic components, $\hat{E}_{i,t}^Y$ and $\hat{E}_{i,t}^X$, they suggest to use the panel no cointegration tests of Pedroni (1999). In particular, we consider the pooled and group mean coefficient test, Z_ρ^p and Z_ρ^{gm} , as well as the pooled and group mean t -test, Z_t^p and Z_t^{gm} . Gengenbach et al. (2006) propose the Johansen trace test to test for cointegration between the estimated common factors \hat{F}_t^Y and \hat{F}_t^X . Denote the statistic as L_{trace}^F . As discussed in Gengenbach et al. (2006), rejection of the null hypothesis of no cointegration for both the idiosyncratic component and the common factor is a necessary but not a sufficient condition for cointegration between $Y_{i,t}$ and $X_{i,t}$.

Gengenbach et al. (2008) propose tests for panel (no) error correction. Starting from the triangular representation of a cointegrated panel with non-stationary (unobserved) common factors, they derive the following conditional error correction model (ECM).

$$\Delta Y_{i,t} = \alpha_i Y_{i,t-1} + \gamma_{1i} X_{i,t-1} + \gamma_{2i} F_{t-1} + \sum_{s=1}^{p_i} \pi_{1is} \Delta Y_{i,t-s} + \sum_{s=0}^{p_i} \pi_{2is} \Delta X_{i,t-s} + \sum_{s=0}^{p_i} \pi_{3is} \Delta F_{t-s} + \varepsilon_{i,t}, \quad (13)$$

where $\gamma_{1i} = -\alpha_i \theta_{1i}$ and $\gamma_{2i} = -\alpha_i \theta_{2i}$ such that $(1, -\theta'_{1i}, -\theta'_{2i})'$ is the cointegrating vector, $X_{i,t}$ is a $m \times 1$ vector of idiosyncratic weakly exogenous variables, F_t is a $k \times 1$ vector of possibly unobserved, strongly exogenous common factors and $\varepsilon_{i,t}$ is an *i.i.d.* error term. Gengenbach et al. (2008) consider 2 tests for the individual specific null hypothesis of no error correction, namely the t -statistic for $H_{0i}^t : \alpha_i = 0, \tau_{\alpha_i}$, and the Wald test for $H_{0i}^w : \delta_i = 0, w_{\delta_i}$, where $\delta_i = (\alpha_i, \gamma'_{1i}, \gamma'_{2i})'$. Following the ideas of Pesaran (2007), they propose to augment the conditional ECM (13) with cross-sectional averages of $Y_{i,t-1}$ and $X_{i,t-1}$ as well as the contemporaneous and lagged averages of $\Delta Y_{i,t}$ and $\Delta X_{i,t}$. The panel test statistics for the null hypothesis of no error correction for all i are given by the (truncated) averages of the individual specific statistics, denoted as $\bar{\tau}_{\alpha_i}^*$ and $\bar{w}_{\delta_i}^*$ for the t and Wald test, respectively. Gengenbach et al. (2008) provide critical values for both test statistics obtained via stochastic simulations.

We test for cointegration between $TRADE_{ijt}$, GDP_{ijt} and $GDPCAP_{ijt}$ using both test procedures outlined above. Using the BIC_3 criterion of Bai and Ng (2002)⁵ we find one common factor in $TRADE_{ijt}$ and three in the joint panel of GDP_{ijt} and $GDPCAP_{ijt}$. The coefficient based tests Z_ρ^p and Z_ρ^{gm} cannot reject the null of no cointegration between the

⁵The results are qualitatively robust to using different criterions.

estimated common factors. However, both t -tests, Z_t^p and Z_t^{gm} , can reject the null. The Johansen trace test, L_{trace}^F , applied to the 4 estimated common factors finds a single cointegrating relationship⁶. For the tests of panel no error correction, $\bar{\tau}_{\alpha_i}^*$ and $\bar{w}_{\delta_i}^*$, allowing for a constant and linear trend in the ECM, we select the lag length p_i using the Akaike information criterion with $pmax = 4$. The critical value for $\bar{\tau}_{\alpha_i}^*$ is -4.040 at 5% significance while the corresponding critical value for $\bar{w}_{\delta_i}^*$ is 28.203 (see Gengenbach et al., 2008, Tables 3 and 4). With statistics of -4.958 and 62.21 for the t and Wald test respectively we can reject the null of no error correct. As only the coefficient based tests for the estimated idiosyncratic components cannot reject the null of no cointegration, there is evidence that $TRADE_{ijt}$, GDP_{ijt} and $GDPCAP_{ijt}$ are cointegrated.

4 Estimation of the gravity equation

In the previous section we have found evidence that the variables entering the gravity equation are non-stationary and cointegrated. Therefore, equation (1) describes a long-run equilibrium relationship between $TRADE_{ijt}$, GDP_{ijt} and $GDPCAP_{ijt}$. In this section, we use the CUP estimator of Bai et al. (2009) and the CCEP estimator of Pesaran (2006) to obtain estimates of the parameters of the static long-run model given in (1). Furthermore, we estimate a dynamic error correction model with a CCEP estimator. All considered estimators allow for a heterogenous effect of the common factors. Furthermore, we allow for heterogenous short-run dynamics when estimating the ECM.

Bai et al. (2009) consider the problem of estimating the cointegrating vector in a cointegrated panel data model with non stationary common factors. They consider the following model,

$$Y_{i,t} = \beta X_{i,t} + \gamma_i F_t + e_{i,t}, \quad (14)$$

where F_t is a k vector of common factors, γ_i is the corresponding vector of factor loadings and e_{it} is an idiosyncratic error term. Bai and Kao (2006) propose a 2-stage fully modified (FM) estimator of β in the case of stationary F_t . However, if F_t is non-stationary the least-squares (LS) estimator of β is inconsistent. Bai et al. (2009) propose a bias corrected (BC) and fully modified (FM) estimator for β for the case of observed F_t . However, those are infeasible in the case of unobserved common factors. The proposed solution is an iterative procedure where F_t is estimated given an estimate of β and then β is re-estimated given the estimate of F_t . The objective function of the procedure is

$$S_{NT}(\beta, F, \Gamma) = \sum_{i=1}^N \sum_{t=1}^T (Y_{i,t} - \beta X_{i,t} - \gamma_i F_t)^2, \quad (15)$$

⁶We also select the cointegration rank using an information criteria advanced by Aznar and Salvador (2002) which finds 3 cointegrating relation ships.

which is minimized subject to the constraints $T^{-2} \sum_{t=1}^T F_t F_t' = I_k$ and $\Gamma' \Gamma$ is positive definite, where $\Gamma = (\gamma'_1, \dots, \gamma'_N)'$.

Now, given F_t the LS estimator of β is given by

$$\hat{\beta} = \left(\sum_{i=1}^N \sum_{t=1}^T Q_F X_{it} (Q_F X_{it})' \right)^{-1} \left(\sum_{i=1}^N \sum_{t=1}^T Q_F X_{it} Q_F Y_{it} \right), \quad (16)$$

where Q_F is the OLS projection error operator such that

$$Q_F X_{i,t} = X_{i,t} - \sum_{t=1}^T X_{i,t} F_t' \left(\sum_{t=1}^T F_t F_t' \right)^{-1} F_t.$$

Define

$$W_{i,t} = Y_{i,t} - \beta X_{i,t},$$

and $W_i = (W_{i,1}, \dots, W_{i,T})'$. The estimator of F_t given $\hat{\beta}$ is given by the k largest eigenvectors of the matrix $(NT^2)^{-1} \sum_{i=1}^N W_i W_i'$. Concentrating out Γ , the objective function can be rewritten as

$$S_{NT}(\beta, F) = \sum_{i=1}^N \sum_{t=1}^T (Q_F W_{i,t})^2.$$

The continuously updated (CUP) estimator is then given by

$$(\hat{\beta}_{CUP}, \hat{F}_{CUP}) = \arg \min S_{NT}(\beta, F).$$

The procedure outlined above requires that k , the number of common factors, is known. In general that is not the case and k has to be estimated. Bai and Ng (2002) propose several information criteria which can be used to obtain consistent estimates.

We obtain CUP estimates of β_1 and β_2 in (1) after concentrating out the fixed effects. We estimate the number of common factors using the BIC_3 criterion of Bai and Ng (2002) which performs well in empirical studies. With $k_{max} = 4$ we estimate a single common factor, such that $\hat{k} = 1$. We then obtain estimates of the long run parameters β_1 and β_2 and estimates of the coefficient of the two dummy variables $EURO_{ijt}$ and FTA_{ijt} which minimize (15).

Pesaran (2006) proposes a consistent estimator for the slope parameter β_i in a heterogeneous panel data model similar to (14). He allows for both observed and unobserved common factors, D_t and F_t respectively. Furthermore, he assumes that $X_{i,t}$ also permits a common factor structure. In particular, his model is given by the following equations:

$$Y_{i,t} = \beta_i X_{i,t} + \alpha_i D_t + \gamma_i F_t + e_{i,t}, \quad (17)$$

$$X_{i,t} = A_i D_t + \Gamma_i F_t + v_{i,t}, \quad (18)$$

where $e_{i,t}$ and $v_{i,t}$ are idiosyncratic errors. Pesaran (2006) suggests to use the cross-sectional average of $Z_{i,t} = (Y_{i,t}, X'_{i,t})'$ as a proxy for the unobserved common factors. The pooled

estimator for β is now give by

$$\hat{\beta}_{CCEP} = \left(\sum_{i=1}^N \sum_{t=1}^T Q_{D,\bar{Z}} X_{it} (Q_{D,\bar{Z}} X_{i,t})' \right)^{-1} \left(\sum_{i=1}^N \sum_{t=1}^T Q_{D,\bar{Z}} X_{it} Q_{D,\bar{Z}} Y_{i,t} \right). \quad (19)$$

Pesaran (2006) shows that β_{CCEP} is consistent for the mean of β_i . However, he only considers weakly stationary variables. Kapetanios et al. (2008) expand the analysis to allow for non-stationary common factors and show that the CCEP estimator remains consistent.

Augmenting the gravity equation (1) with cross-sectional averages of $TRADE_{ijt}$, GDP_{ijt} and $GDPCAP_{ijt}$ we obtain CCEP estimates of β_1 and β_2 as well as of δ_1 and δ_2 .

Furthermore, we estimate a dynamic ECM as given in (13). We obtain CCEP-type estimates by including cross-sectional averages the lagged level and (lagged) first-differences of $TRADE_{ijt}$, GDP_{ijt} and $GDPCAP_{ijt}$ as a proxy for the common factors in the regression. Allowing for a maximum lag length of $p_{max} = 6$ we select an appropriate lag length of $\hat{p} = 4$ using the *BIC*.

Table 3 reports the obtained parameter estimates as well as the results obtained by Bun and Klaassen (2007, 2002)⁷ for direct comparison. For the ECM estimates results are sensitive to the specified lag length. We report findings for the estimated lag length $\hat{p} = 4$ as well as for estimates obtained without allowing for additional short-run dynamics in the model ($p = 0$). Similarly to the findings of Bun and Klaassen (2007), the CUP estimate and the CCEP-ECM estimate for the case without short-run dynamics observed a strong drop in the estimated coefficients when allowing for country pair specific trends. Without trends, the CUP estimate of the Euro effect on trade is about 52% and the CCEP-ECM estimate even 62%. However, these estimates are reduced to 7.8% and 3.4%, respectively, in the trend case. The CCEP estimator for the static model finds a Euro effect on bilateral trade of about 4.4% when not allowing for trends which is reduced to only 0.6% in the trend case. The CCEP-ECM estimate in the model with short-run dynamics is even negative in the no trend case with an estimated long-run effect of about -3.7%. When allowing for trends, the effect is estimated at about 1.5%.

Estimates for β_1 and β_2 also vary between estimators and models. While the estimates of β_1 all have the expected sign, the static CCEP estimate in the no trend case and the CCEP-ECM estimate with short-run dynamics and trend are very small at 0.014 and 0.027, respectively. For these two estimators we observe a stronger positive effect of GDP per capita than GDP on trade, with estimates of β_2 at 0.354 and 0.731, respectively. In 3 cases we obtain negative estimates of β_2 , namely for the CUP estimator without trends and for the CCEP-ECM estimators without short-run dynamics. In those cases the estimate of β_2 is smaller than the estimate of β_1 in absolute value. Our estimates of the trade effect of a free

⁷Bun and Klaassen (2002) use a different data set. However, we include their results as a comparison for the dynamic model.

Table 3: Parameter estimates for static and dynamic models

Static model estimates								
	Bun and Klaassen (2007)				current study			
	LSDV		DOLS		CCEP		CUP	
	no trends	trends	no trends	trends	no trends	trends	no trends	trends
$\hat{\beta}_1$	1.410	0.700	0.590	0.940	0.014	0.334	1.351	0.359
$\hat{\beta}_2$	-0.680	-0.230	0.200	-0.490	0.354	0.091	-0.710	0.179
$\hat{\delta}_1$	0.410	0.032	0.374	0.034	0.043	0.006	0.417	0.075
$\hat{\delta}_2$	0.410	0.060	0.380	0.050	0.033	0.011	0.418	0.016
Dynamic model estimates								
	Bun and Klaassen (2002)				CCEP for ECM			
	LSDV for ADL				$p = 0$		$\hat{p} = 4$	
	no trends	trends			no trends	trends	no trends	trends
$\hat{\alpha}$	-	-			-0.407	-0.487	-0.782	-0.785
$\hat{\gamma}_1$	-	-			0.303	0.340	0.257	0.021
$\hat{\gamma}_2$	-	-			-0.238	-0.232	0.170	0.574
$\hat{\delta}_1$	0.040	-			0.196	0.016	-0.030	0.012
$\hat{\delta}_2$	0.080	-			0.025	0.008	0.020	0.008
Long-run estimates								
	no trends	trends			no trends	trends	no trends	trends
$\hat{\beta}_1$	-	-			0.744	0.698	0.329	0.027
$\hat{\beta}_2$	-	-			-0.585	-0.476	0.217	0.731
$\hat{\delta}_1$	0.330	-			0.482	0.033	-0.038	0.015
$\hat{\delta}_2$	0.710	-			0.061	0.016	0.026	0.010

Notes: “No trends” indicates that τ_{ij} is set to 0. LSDV gives results for LSDV-type estimates from Bun and Klaassen (2007, Table 2, p. 480). DOLS give the estimates from Bun and Klaassen (2007, Table 5, p. 491). Dynamic model estimates for a stationary ADL are taken from Bun and Klaassen (2002, Table 1, p. 11), where results are reported for the European FTA dummy.

Results for Bai, Kao, and Ng’s CUP estimator are obtained with $\hat{k} = 1$ common factor, as selected by BIC_3 . For the ECM a lag length of $\hat{p} = 4$ is selected by the BIC.

trade area are in general smaller than previously reported estimates. We find a positive effect between 1% and 6.3%. Only the CUP estimator in the no trend case finds a strong effect of about 52%.

5 Conclusion

Using the data of Bun and Klaassen (2007) we have estimated the Euro effect on bilateral trade using a cointegrated panel data model. Bilateral trade data is by construction strongly cross-sectionally correlated and we have allowed for persistent cross-sectional dependencies by allowing for (unobserved) common factors. Using several panel unit root tests, we have

found strong evidence that the variables entering the gravity equation are non-stationary. However, $TRADE_{ijt}$, GDP_{ijt} and $GDPCAP_{ijt}$ seem to be cointegrated as indicated by the panel cointegration test of Gengenbach et al. (2006) and the panel error correction tests of Gengenbach et al. (2008). Using the CCEP estimator of Pesaran (2006) and the CUP estimator of Bai et al. (2009) we obtain estimates of the parameters of the static long-run model. We also obtain CCEP-type estimates for the parameter of a dynamic ECM. Our parameter estimates vary between models and estimators but seem to confirm the findings of Bun and Klaassen (2007) of a smaller Euro effect than previously estimated. Only the CUP and CCEP-ECM estimator find strong effects of the Euro on trade when not accounting for country pair specific trends.

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