Professor Dumitru MIRON, PhD
Email: dumitru.miron@ase.rei.ro
Professor Paul MICLĂUS, PhD
Email: pmiclaus@gmail.com
Danusia VAMVU, PhD student
Email: danusia_vamvu@yahoo.com
Faculty of Economics and International Affairs
The Bucharest Academy of Economic Studies

THE EFFECT OF CURRENCY UNIONS ON TRADE : TOWARDS BETTER ESTIMATION TECHNIQUES

Abstract: *A gravity model is used to assess the effects of currency unions on international trade. The panel data set used includes bilateral observations for eight years spanning 1980 through 2013 for 186 countries. In this data set, there are 477 observations in which both countries use the same currency. We find a decreasing positive effect of currency union on international trade when using different estimation techniques. Even though the results are much more reliable than those obtained by Rose (2000), we still find a series of issues and have reservations in association with the related techniques and the results should be treated accordingly.*

Keywords: *Rose effect, common currency, trade, gravity model, monetary union.*

JEL Classification: F33

1. Introduction

The quest to separate out the effect of currency unions on trade (henceforth the Rose effect) will take its point of departure in the first paper that was ever written on the topic, Rose (2000). We should immediately point out that the econometric techniques as well as the data itself used in the paper can easily be shown to contain a large number of deficiencies which distorts the estimates but reproducing the basic first findings is useful for illustrative purposes and it enables us to better identify the sources of the likely biases. The paper exploits a large dataset of 33,903 bilateral trade observations spanning five different years (1970, 1975, 1980, 1985 and 1990) to estimate a conventional gravity equation. A dummy for currency union status which is equal to one if the country pair is using the same currency is included and the size of the Rose effect as well as statistical significance can then be assessed. The benchmark regression in Rose (2000) takes the following form:

(1) $\ln(RV_{od})=a_0+\beta_1\ln(RY_oRY_d)+\beta_2\ln(Distance_{od})+\beta_3(CU_{od})$ + controls

where RV_{od} is the real value of bilateraltrade, the RYs are real GDPs of the origin and destination countries, and CU is a dummywhichis 1 when the destination and the origin country share the samecurrency. The coefficient of interest to Rose(2000) was, obviously, β_3 (the effect of a currency union on trade, alsoknown as "the Rose effect").

The coefficient of interest to us is, obviously, β_3 - the coefficient on the currency union (CU) dummy, or the Rose effect. Prior to Rose (2000) the literature mainly assumed that joining a currency union would produce trade effects similar in magnitude to reducing bilateral exchange rate volatility to zero (e.g. Frankel and Rose 1998). Unfortunately, the results in this area were disappointing in the sense that bilateral exchange rate volatility, if it has an effect at all, is extremely difficult to measure even with advanced time-series data, and hence, increasing trade cannot be used as a primary motive for two countries to adopt a common currency. Euro-sceptics have argued that since exchange rate volatility between the Euro-zone countries was low prior to adopting the Euro, and since this exchange rate uncertainty could be easily hedged through forward contracts, trade would only increase by a small and insignificant amount by adopting the Euro. Proponents, on the other hand, argue that a currency union is a much deeper form of economic integration and is likely to produce trade-creating effects of a greater magnitude due to the empirically observed home-bias in trade. McCallum (1995) estimates this home bias to be more than twenty to one. He finds that two Canadian provinces trade twenty times as much as a comparable American state/Canadian province pair.

The advantage of (1) is that it includes a measure of bilateral exchange rate volatility (*EXRATEVOL*) as well as a currency union dummy[1], and hence, we are then able to separate out their individual contributions. We note that the point estimates of the Rose effect get higher over the years in the sample: it varies from a minimum of 0.87 in 1970 to a maximum of 1.51 in 1990. The pooled sample estimate (which should give us a long-term measure of the Rose effect) is 1.21. This coefficient estimate implies, *caeteris paribus*, that countries that use a common currency trade three times as much as countries using different currencies ($e^{1.21}\approx 3.35$). In other words, a currency union boosts trade by 235%.

[1]Bilateral exchange rate volatility is measured as the standard deviation of the first differences of the natural log of the bilateral exchange rate in the five years preceding year *t* (Rose 2000 experiments with alternative measures of exchange rate volatility).

2. Literature review or the Rose vine

2.1. Problems that demand addressing

We now have to decide whether we want to accept this point estimate or whether we should stop and smell the roses. From an intuitive point of view does it make sense – after controlling for a host of factors related to trade – that the trade-creating effect of common currency is more than 200%? I would say no, and we now have to find out what is going on.

Knowing that the large Rose effect would be met by objection, Rose (2000) himself carried out an immense amount of sensitivity checks (most of which cannot be summarised here due to lack of space) to show that his results remained robust to a number of specifications.

The source of bias that Rose (2000) is concerned with is reverse causality (countries that trade more are more likely to form monetary arrangements to reduce transaction costs). To overcome this problem Rose (2000) experiments with IV estimation. Three terms involving inflation rates are used as instruments for the currency union dummy and exchange rate volatility; i) the product of the two countries' inflation rates, ii) the sum of the inflation rates and, iii) the absolute difference between the inflation rates[2]. However, since the CU dummy now becomes "wildly and implausibly bigger" (Rose 2000 p. 22), these IV estimates are difficult to trust. Although the standard Hausman tests in Rose's data cannot reject the null hypothesis of exogeneity of the CU dummy[3], there are good theoretical reasons to suspect simultaneity biases, and also that these biases are higher for smaller, poorer and more open economies. A political economy framework suggests a theory: a depreciation of a country's currency against its major trading partner will lead to an increase in exports and a decrease in imports and vice-versa in case of an appreciation. Since importers and exporters dominate politics in small and poor countries (Baldwin 2006 p. 32) and since losers (either the importers or exporters) usually lobby harder, political pressures will ensure that the volatility of the exchange rate is stable against a country's major trading partner and in extreme cases this will lead to adopting a common currency. The reverse causality problem could thus be expected to be more serious in Rose's dataset.

Baldwin (2006) mathematically derives the likely endogeneity issues. He identifies a term which he refers to as the "relative-prices-matter" term which is usually called the multilateral trade resistance term in the literature. He argues that since this term is missing in the conventional gravity equation it is likely that the CU dummy picks up this effect[4].

[2] All measures are calculated in the five years preceding year *t*.

[3] Rose (2000) points out that these tests should be taken with a grain of salt, since it is difficult to find good instruments.

[4] It turns out that the "relative-prices-matter" term actually contains the CU dummy, and since it is omitted it will be left in the error term and produce a bias. See Baldwin (2006) pp. 5-20 for a mathematical derivation of this.

The most easily understood example of this is the distance between countries which is an inadequate measure of the true trade costs. The distance between Australia and New Zealand is 2312 kilometres in my dataset but a flight from Melbourne to Wellington is only three hours and twenty minutes. Compared with flights to other developed nations' major cities (9 hours to Tokyo, 16 hours to Los Angeles and 23 hours to London[5]) the Australia-New Zealand trade costs are not significantly high relative to the trade costs faced by exporters from the rest of the world. The problem in Europe might work in reverse to that of Australia and New Zealand. All the developed countries in Europe are very close but will not experience significantly lower trade costs than between Australia and New Zealand.

The coefficient of the log of distance in gravity equations is typically around -0.7 for European samples (see Baldwin 2006 p.14) but for my sample and also in Rose's (2000) dataset the log of distance coefficient is slightly lower than -1 which means that the distance coefficient will explain more of the variation in the dependent variable and, hence, it will incorrectly produce more upward or downward biases (in the case of Australia-New Zealand the gravity equation under predicts the trade costs whereas the reverse is the case in Europe).

Further problems arise for the developing world. Because infrastructure is less developed, distance is a completely different concept. Trade costs between New Delhi and Calcutta are immensely high compared to two developed-country cites of similar distance and size. Also, distance across land is different than distance across oceans.

2.2. Different techniques

The problem with the conventional gravity equation is that regardless of how many explanatory variables we add to the regression we cannot possibly account for the numerous unobserved fixed effects that are responsible for both the level of trade between two nations and whether two nations share a common currency. The unique historical relationship between the African CFA countries is very different from that of, say, Australia and its territories and dependencies. In order to identify the real source of correlation between currency-union status and trade, one should essentially be an expert in a particular currency area. In 1979, Ireland broke its 1:1 peg with the British pound.

Taking Rose's (2000) estimate literally this should have decreased the Anglo-Irish trade to one third of what it was prior to the break-up (holding a host of other factors constant). This clearly did not occur. The UK's share of Irish trade was roughly 50 % in 1979 and five years later it dropped to roughly 40 % (see Thom and Walsh 2002). However, there has been a steady decline in Anglo-Irish trade since 1924, and the explanation for this decline is more likely to reflect the fact that Ireland became less dependent on the UK over this period. The reduced

[5]Sourse: www.qantas.com.au.

dependency on the British market was driven by factors that are unobservable to the conventional gravity model.

Thom and Walsh (2002) find no statistically significant decline in trade from neither time-series nor panel data following the break-up. To account for the problem of uniqueness of each individual currency union we might want to apply different estimation techniques to Rose's (2000) data. The paper, to my knowledge, that first addresses this issue is Rose and Van Wincoop (2001). The unobservable characteristics (omitted variables) are dealt with by including origin-nation and destination-nation dummies (i.e. a rose for every country) on a dataset for the time period 1970-1990 with controls similar to those in Rose (2000). The result is a smaller Rose effect. The coefficients on the CU dummy are without and with country-dummies, respectively, 1.38 and 0.86[6]translating into percentage Rose effects of roughly 297% and 136%, respectively. In other words, a large drop! This result is more promising than the original estimate because it controls for country-specific omitted variables in the cross-sectional regression.

The problem is, however, that many of the omitted variables might be time-varying and including time-invariant country-dummies does not take care of that problem. The "relative-prices-matter" term may at first appear to be time-invariant but the concept of trade costs may very well change over time. Faster internet connections and more virtual space to organise business meetings may reduce bilateral trade costs. Moreover, the reduced dependency of Ireland on the British market may also be driven by time-varying factors[7].

Other papers have included country-pair dummies. Rose (2001) explains that the reason for not including pair-specific dummies in his original paper is that this procedure wipes out all cross-sectional variation leaving only time-series variation to explain the Rose effect. This is problematic as the data in Rose (2000) covers the period 1970-1990 – a period where there was hardly any variation in the CU dummy. When Rose (2001) includes pair-specific dummies on the original data the Rose effect drops to -0.38 with a standard error of 0.67.

Independently of Rose (2001), Pakko and Wall (2001) collect a dataset for the same period. As in my own dataset, Pakko and Wall (2001) maintain direction-specific country pairs and, therefore, when including pair-dummies they are able to capture any asymmetries of the Rose effect. Pakko and Wall (2001) obtain a coefficient on the CU dummy of -0.38 with a standard error of 0.529. Although the actual point estimates of the Rose effect are negative, a positive Rose effect can be obtained within one standard error of both estimates. But the reliability of the actual estimates is highly questionable.

With hardly any variation in the CU dummy, the results come from variation of only a few observations of entry or exit into currency union (see the appendix) so the estimates are not very informative about the actual Rose effect.

[6]The standard errors of the two estimates are 0.19.

[7]However, as the CU dummy did not vary a great deal over this time period it might not be a big problem.

Whereas Rose (2001) rejects the estimate on those grounds, Pakko and Wall (2001) conclude "while it may well be true that the statistical insignificance of the common currency dummy should not be taken to mean that the effect is not positive, this misses the point.

A comparison of the two sets of results suggests that pooled cross-section estimates are not reliable because they are biased by the mismeasurement of trading-pair specific variables. This is evident in the dramatically different coefficients on the GDP and per capita GDP variables[8] that are found using the two methods. In other words, the restrictions necessary to obtain the pooled cross-section specification from the fixed-effects specification are rejected, indicating that the fixed-effects specification is preferred". (Pakko and Wall 2001 p. 43).

With this in mind Glick and Rose (2002) set out to collect a dataset that covers the period 1948-1997 with more variation in the CU dummy. The pooled cross-section estimate of the conventional gravity equation with controls similar to Rose (2000) turns out to be 1.3 (267 %) whereas using country-pair fixed effects and other time-varying controls reduce this estimate to 0.65 (92 %).

It was argued above that throwing in time-invariant country dummies as in Rose and Van Wincoop (2001) does not eliminate biases stemming from variation over time of the "relative-prices-matter" term and other time-varying factors that are correlated with the CU dummy. Since country-pair dummies are also time-invariant they would do no better. In other words, there is no guarantee that the pair-specific factors that are captured by the dummies remain constant over time.

3. Research methodology or the updated garden of roses

As it would be interesting to see the results of applying all of the above techniques to my own data (which covers a different time period), these are presented in ***Table 1*** from the 4th section of this article.

We collected an updated dataset for eight years spanning through 1980 – 2013 with almost the same variables as in Rose (2000). The full updated sample consists of 49,709observations on bilateral trade (the dependent variable) of 186 countries (I lack GDP data for a number of the countries, however, just as in Rose 2000) for the years 1980, 1985, 1990, 1995, 2000, 2005, 2010 and 2013.

Note that after 1999, extra CU-pairs corresponding to the Euro-zone were added to the model. This brings the number of CU pairs to 477.

Also note that my dataset is not balanced including fewer bilateral trade observations prior to 1999 due to the lack of information on export and import flows mainly for the underdeveloped and developing countries.

We use an augmented gravity model to estimate the effects of currency unions on trade. The model is "augmented" in the sense that the standard gravity model only includes the natural logarithms of income and distance variables. In

[8] The GDP and GDP per capita coefficients are, respectively, 1.34 and -0.151 using country-pair fixed effects on the data in Pakko and Wall (2001).

order to account for as many other factors as possible, my equation adds a host of extra conditioning variables but, unfortunately, it does not include all the important monetary variables:

(1) $\ln(TE_{ijt}) = \beta_0 + \beta_1 lnPrGDP_{ijt} + \beta_2 lnPrGDPPC_{ijt} + \beta_3 CU_t + \beta_4 FTA_t + \beta_5 border + \beta_6 lndistance + \beta_7 com\_lang + \beta_8 com\_col + \beta_9 colonial + \beta_{10} landlocked + \beta_{11} oneisland + \beta_{12} twoisland + \beta_{13} exrate\_vol_t + \beta_{14} yeardummies$

where *i* and *j* denotes exporting and importing countries respectively, *t* denotes time and the variables are defined as follows:

TE_{ijt} denotes nominal exports from country *i* to country *j* measured in million US dollars deflated by US consumer price index (main sources for export flows: World Bank and International Monetary Fund, Direction of Trade Statistics; main sources for CPI: International Financial Statistics, IMF);
$PrGDP_{ijt}$ denotes the product of the real GDP of country *i* and country *j* in constant 2000 million US dollars (main source: IMF);
$PrGDPPC_{ijt}$ denotes the product of the real GDP per capita of country *i* and country *j* in constant US dollars (main source: IMF);
CU_t denotes a binary variable equal to 1 if country *i* and country *j* use the same currency (main sources: 1980-1990 Rose(2000) 1991-2013 own research, mainly using IMF);
FTA_t denotes a binary variable equal to 1 if country *i* and country *j* have a bilateral free trade agreement (main source: World Trade Organization);
Border denotes a binary variable which equals 1 if country *i* and country *j* share a land border (own research, mainly using Google maps);
Distance denotes the distance measured in kilometers between country *i* and country *j* (main source: World Bank);
Com_lang denotes a binary variable which equals 1 if country *i* and country *j* have a common official language (main source: World Bank);
Com_col denotes a binary variable which equals 1 if country *i* and country *j* were colonies after 1945 with the same colonizer (main source: Glick and Rose(2002));
Colonial denotes a binary variable whichequals 1 if country *i* colonized country *j* or vice versa (main source: Glick and Rose(2002));
Oneisland denotes a binary variable which equals 1 if any of the country *i* or country *j* are islands and 0 otherwise (own research);
Twoisland denotes a binary variable which equals 1 if both countries in a pair are islands and 0 otherwise (own research);
Landlocked denotes a variable whichequals 2 if both countries are landlocked, 1 if any of the country *i* or country *j* are landlocked and 0 if none of them are landlocked (own research);
Exrate_vol denotes the exchange rate volatility between country *i* and country *j* measured as the standard deviation of the first differences of the natural log of the

bilateral exchange rate in the five years preceding year *t*, as defined in Rose (2000); (main source: International Financial Statistics, IMF).

4. Results

Let us briefly sum up our findings so far. We collected a large updated dataset similar to Rose's (2000) but the data problems that were present in Rose are certainly also present in our dataset for the period 1980-1995. The common currency areas mainly consist of small, poor and open countries, each of their individual historical reasons to form a currency union are unique, and the self-selection process into these monetary unions was largely non-random over the period. The last four years of my data include many more observations from the Euro-zone, and hence, countries that were more randomly selected for currency union at least if we believe Andrew Rose: "…few contemporary commentators believe that EMU was mostly pursued for economic motives." (Rose 2000 p. 22).

However, even if this is true, the Euro-zone is a "young" currency area and we cannot be sure that the long-run effect shows up in the data at this early stage.

From ***Table 1***, we note that using country dummies shrinks the Rose effect from 0.745 to 0.504 (or from 110% to 65%). The time variation of the CU dummy comes from the Euro-zone and the few countries that adopted the US dollar so the estimate of the Rose effect should, in theory, come from these observations. The fourth column estimates a regression based on country-pair fixed effects. This means a regression of (2) but excluding all time-invariant variables such as border, distance and common language. Ideally we should have added dummies for each country pair but since there are so many country pairs (more than 14,000) we use country-pair fixed effects which essentially gives the same results (except for a lower R^2 because including dummies only increases R^2 because there are more explanatory variables). The estimate of the Rose effect remains positive but very small and insignificant, 0.077 (nearly 8%). This estimate relies on slightly more observations of switches in currency union status than the estimates of Rose (2001) and Pakko and Wall (2001) but my result here is still hard to trust, especially due to the persisting issues that are associated with our techniques.

TABLE 1 – Various estimation techniques applied to our dataset

	Pooled OLS	**Pooled OLS with country dummies**	**Country-pair fixed effects**
GDP	**0.93*** (0.004)**	**0.901*** (0.094)**	**1.455*** (0.065)**
GDP/capita	**0.205*** (0.009)**	**0.248** (0.085)**	**-0.082 (0.059)**
CU	**0.745*** (0.08)**	**0.504*** (0.09)**	**0.074 (0.099)**

FTA	**0.443*** (0.029)**	**0.36*** (0.033)**	**0.246*** (0.038)**
Exratevol	**-0.01*** (0.001)**	**-0.007*** (0.001)**	**-0.006*** (0.001)**
Border	**0.825*** (0.061)**	**0.509*** (0.061)**	
Distance	**-1.127*** (0.013)**	**-1.548*** (0.015)**	
Com_lang	**0.557*** (0.033)**	**0.817*** (0.034)**	
Com_col	**1.165*** (0.046)**	**0.961*** (0.044)**	**1.403** (0.626)**
Colony	**1.3*** (0.06)**	**1.155*** (0.052)**	**-0.202 (0.598)**
Percentage Rose effect	**110%**	**66%**	**8%**
Number of Observations	**49,709**	**49,709**	**49,709**
Number of CU=1 observations (% of total)	**0.96%**	**0.96%**	**0.96%**
R^2	**0.64**	**0.77**	**0.53**
RMSE	**2.16**	**1.74**	**N/A**

*** significant at 10%**
**** significant at 5%**
*****significant at 1%**
Note: Robust standard errors in parentheses
Years: 1980, 1985, 1990, 1995, 2000,2005, 2010 and 2013

Source: own calculations using data collected as mentioned in section 3.

5. Conclusions

The unanimous conclusion to draw from all of the alternative estimations that were applied on Rose's (2000) and other datasets is a shrinking of the Rose effect over the years.

As endogeneity is highly likely in our data, it appears so far that the most reliable estimate of a Rose effect is that obtained using the above presented techniques. Recall that we rely on a long time period and use country-pair dummies to strip off the correlation between the CU dummy and omitted variables.

But since we are concerned that time-varying factors are correlated with the CU dummy there are still reasons to be worried.

Moreover, worryingly, the percentage Rose effects obtained from the point estimates of the cross-sectional regressions in the Glick and Rose (2002) dataset, increases over time. While the Rose effects prior to 1985 appear to vary between 100 % and 240 % they vary between 350 % and 1000 %[9] in 1985-1995. In other words, the effects are implausibly large towards the end of the period. Baldwin (2006) suggests that the group of currency unions in 1950 was relatively random whereas currency union leavers over the period were surely not random. Small, poor and open economies such as Kiribati could not afford to leave the currency union arrangement with Australia whereas Algeria could afford to drop out of its currency union with France and other African nations.

Hence, the degree of non-random selection into or out of currency unions appears to get more severe towards the end of the 20th century which is bad news for Rose (2000) and ourselves.

The discussion of non-random selection into currency unions may also motive us to think about why the results of our cross-sectional and pooled cross-sectional estimates indicated that the trade-creating effects of currency union are so much greater than the effects of bilateral free trade agreements and reducing bilateral exchange rate volatility: the two latter trade arrangements emerge out of a somewhat more random process.

Moreover, since bilateral free trade and exchange rate arrangements occur more frequently, there are many more observations of relevance and sample selection biases are less likely (the low percentage of currency union pairs in ***Table 1***).

All the above mentioned issues make us reiterate the fact that effort should still be devoted to overcoming these problems and finding better estimation techniques.

The real answer to the question “How much will the Euro boost trade?” will probably not appear in the aggregate trade data for many years. But while we wait for this to appear, focus should probably be diverted to carrying out surveys that aim to ask business leaders, who are using the Euro, how exactly the Euro has changed the way they do business. This might give us a better idea of what to expect of future currency unions.

[9]The Rose effects obtained in our dataset for these years are also rather large.

REFERENCES

[1] **Alesina, A. and Barro, R. (2000)*, Currency Unions;*** *NBER Working Paper No. W7927;*
[2] **Baldwin, R. (2006)*, The Euro's Trade Effects;*** *European Central Bank Working Paper;*
[3] **Baldwin, R., Skudelny, F. and Taglioni, D. (2005), *Trade Effects of the Euro: Evidence from Sectoral Data;*** *Working paper series 446, European Central Bank;*
[4] **Berger, H. and Nitsch, V. (2005), *Zooming Out:The Trade Effect of the Euro in Historical Perspective;*** *CESIFO Working Paper No.* 1435;
[5] **Flam, H. and Nordstrom, H. (2003)*, Trade Volume Effects of the Euro: Aggregate and Sector Estimates*;**
[6] **Frankel, Jeffrey A. and Rose, A. (1998)*, The Endogeneity of the Optimum Currency Area Criteria*;** *Economic Journal*;
[7] **Glick, R. and Rose, A. (2001)*, Does a Currency Union Affect Trade? The Time Series Evidence;*** *NBER Working Paper No. 8396*;
[8] **McCallum, J. (1995), *National Borders Matter: Canada-US Regional Trade Patterns;*** *American Economic Review 85-3, 615-623;*
[9] **Micco, A., Stein, E. and Ordonez, G. (2003), *The Currency Union Effect on Trade: Early Evidence from EMU"*;** *Inter-American Development Bank;*
[10] **Nitsch, V. (2004), *Have a Break, have a…National Currency: when do monetary unions fall apart?;*** *CESIFO Working Paper No.* 1113;
[11] **Nitsch, V. (2005), *Currency Unions and Trade;*** *Globalisation of Capital Markets and Monetary Policy, Edited by Jens Hölscher and Horst Tomann, Palgrave MacMillan, 2005;*
[12] **Pakko, M. and Wall, H. (2001)*, Reconsidering the Trade-Creating Effects of a Currency Union***; *FBR St. Louis Review* 83-5, 37-45;
[13] **Persson, T. (2001), *Currency Unions and Trade: How Large is the Treatment Effect?*;** *Economic Policy 33,435-448;*
[14] **Rose, A. (2000), *One Money, One Market: Estimating the Effect of Common Currencies on Trade;*** *Economic Policy, Vol. 15 Issue 30 pp.7-46;*
[15] **Rose, A. and Van Wincoop, E (2001), *National Money as a Barrier to International Trade*;** *http://faculty.haas.berkeley.edu/arose/;*
[16] **Rose, A. (2001); *Currency Unions and Trade: the Effect is Large;*** *Economic Policy 33, 449-461;*
[17] **Thom, R. and B. Walsh (2002), *The Effect of a Currency Union on Trade: Lessons from the Irish Experience;*** *European Economic Review 46, 1111-23.*