MONETARY UNIONS AND TRADE :

TOWARDS A MORE RELIABLE METHOD OF ESTIMATION

Julie Lochard ♦

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I. Introduction

The Optimum Currency Area (OCA) theory (Mundell, 1961) states that a currency area should increase trade by eliminating transaction costs. This largely shared approach leads to consider a monetary union – i.e. different countries using the same currency – as equivalent to a fixed exchange rate regime. In this context, the paper published by Rose (2000) made a great fuss. Using a gravity model and a large sample including more than 180 countries, Rose highlights that the effect of a currency union may differ from the effect of a peg: while fixing the exchange rate by eliminating its volatility has a negligible impact on the volume of trade, a currency union, on the contrary, enhances trade by a factor of three. This finding, if it is confirmed, has strong implications for further research. It means that the way of understanding currency unions and its impact on trade has to be changed: one cannot simply assume that a currency union just favours trade by reducing transaction costs, since in theory, a fixed exchange rate regime has the same effect.

In the scope of this paper, we will try to check for the robustness of Rose's result, using his own data set. Several authors studied this issue, and many reach the conclusion that Rose's estimate is not reliable because of an omitted variable bias. For instance, Tenreyro (2001) states on page three: "Compatibility in legal systems, cultural links, better infrastructure and tied bilateral transfers, for example, may increase the propensity to form a currency union as well as strengthen trade links between countries. This correlation could lead

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♦ ROSES, University of Paris I Panthéon-Sorbonne. E-mail: jlochard@univ-paris1.fr. For helpful comments I would like to thank Mathilde Maurel, Gérard Duchêne, and José De Sousa. For making his data available, I also thank Andrew K. Rose.

to a positive bias in simple OLS estimates". The major caveat of Rose's estimation is indeed that it does not control for every possible element which is correlated to the currency union variable, while affecting bilateral trade. The only way to control for these factors and hence eliminate a potential bias in the estimation of the currency union variable is to use panel-data techniques. Among these, the *Within* model is generally the most appropriate, since it always provides consistent estimators, even when there is a correlation between the unobservable characteristics and some of the explanatory variables. But the counterpart of this advantage is firstly, that it does not provide fully efficient estimators, and secondly that it does not allow to measure the impact of time-invariant factors. To overcome this problem, Hausman and Taylor (1981) propose a procedure which produces consistent and efficient estimators (including for time-invariant variables).

Re-estimating Rose's gravity equation with this method, we find that the trade-creating effect of a currency union disappears. This result does not mean obviously that a currency union do not have any effect on trade; it only means that it does not have an effect *per se*. For our part, we think that the impact of a currency union on trade is probably more indirect, and thus could not be captured by a simple gravity approach. Further research should then think about the construction of a more general model – integrating other variables, especially in the financial sphere – to capture any indirect effect of monetary arrangements on trade. But this implies first to understand better how a currency union affects trade.

We will proceed as follow. In a second section of this paper, we will present the starting point: Rose (2000)'s method and results. Then, in a third section, we will review the different papers which tried to question this original finding. In the fourth section, we will show that the only way to reach a reliable estimation is to use panel-data techniques, and among these, the method proposed by Hausman and Taylor (1981). The fifth section will be devoted to the presentation of our own results, and in the final section, we will conclude and give some hints for future research.

II. The background: Rose's original estimate

To measure the effect of currency union membership on trade, Rose (2000) estimates a gravity equation. This kind of model is generally used to evaluate and predict the impact of different factors – trade agreements for example – on the volume of bilateral trade. In the more simple specification, the volume of trade between two countries is a function of their level of production and of their relative distance. In general, authors add to these three variables, different dummies to control for factors which can affect trade – like the sharing of a border or the use of a same language. Some authors introduce also a variable of GDP per capita as a proxy for the level of development: as a country becomes more developed, it tends to specialize more and trade more. Rose adds two additional variables: a dummy which indicates if the two countries belong to a currency union, and a variable measuring the volatility of bilateral exchange rates. The estimated equation is:

(1) $$\ln(X_{ij}) = b_0 + b_1 \ln(Y_i Y_j) + b_2 \ln(y_i y_j) + b_3 \ln D_{ij} + b_4 Border_{ij} + b_5 Comlang_{ij} + b_6 Regional_{ij} + b_7 Comctry_{ij} + b_8 Comcol_{ij} + b_9 Colonial_{ij} + gCU_{ij} + dV(e_{ij}) + e_{ij}$$

where X_{ij} is the value of bilateral trade between countries i and j, Y is the real GDP , y is the real GDP per capita, D_{ij} is the distance between i and j, $Border_{ij}$ and $Comlang_{ij}$ are two binary variables which are unity if countries i and j share a land border and a common language, $Regional_{ij}$ is a binary variable which is unity if i and j belong to the same regional trade agreement, $Comctry_{ij}$, $Comcol_{ij}$, and $Colonial_{ij}$ are three other variables of control which are 1 respectively if i and j belong to a same country, if they were colonies with the same colonizer and if i colonized j or *vice versa*. Finally, CU_{ij} is a dummy which is 1 when i and j share the same currency and $V(e_{ij})$ represents the volatility of nominal bilateral exchange rate.

Rose's sample includes more than 180 countries, dependencies, overseas department or colonies located all over the world, and 5 years (1970, 75, 80, 85 and 1990). In this data set, there are 330 observations where two countries trade and belong to the same currency union[1]. Using cross-sectional analysis - a pooled regression with year controls, Rose shows that belonging to a monetary union increases bilateral trade by 235% (= $e^{1.21}$ -1). Hence, countries sharing the same currency trade three times more with each other than with countries using different currencies (see table 1).

[1] The currency unions included here involve for example the following countries: Bahamas/USA, Senegal/other CFA countries, Panama/USA, Guadeloupe/France, Ireland/UK, Luxembourg/Belgium.

TABLE 1 : ROSE'S ORIGINAL ESTIMATE

	OLS ESTIMATION (ROSE, 2000)
Currency Union	1.21*** (.14)
Volatility of exchange rates	-.017*** (.002)
Gdp	.80*** (.01)
Gdp per capita	.66*** (.01)
Distance	-1.09*** (.02)
Regional trade agreement	.99*** (.08)
Common language	.40*** (.04)
Common border	.53*** (.08)
Common colonizer	.63*** (.06)
Common country	1.29*** (.26)
Colonial relationship	2.20*** (.07)
N° of obs.	22 948
R^2	0.63

Notes :
Constant and year-specific fixed effects not reported. Standard errors in parentheses.
* significant at 10%, **significant at 5%, *** significant at 1%.
Source: Rose (2000).

This estimation is very large and Rose recognizes himself that it cannot be easily transposed to out-of-sample countries (like EMU countries for example). A result perhaps more reliable is that the effect of a reduction in exchange rate volatility is radically different from the effect of a common currency. First of all, the common currency coefficient is still significant and large even after introducing a variable of exchange rate volatility, which means that a monetary union has an effect that goes beyond the effect of exchange rate stabilization. This result is a key issue since it questions a relative consensus on the idea that the effect of a common currency on trade is essentially due to an elimination of transaction costs (the OCA approach). But moreover, a monetary union seems to have a quite larger effect on trade, compared to a fixed exchange rate regime. Rose indicates that reducing exchange rate volatility from one standard deviation to zero would increase the value of trade by only 13%, which should be compared to the increase of 235% generated by a common currency. Note that this little effect of volatility on trade is consistent with prior literature. Among the cross-sectional estimations, Frankel and Wei (1993), using data on 63 countries, show that exchange rate variability has only a small effect on trade: if all countries adopt fixed exchange rates, the volume of trade would increase by only 3%. Regarding time-series estimations, the effect estimated is even more weak.

If it is confirmed, this trade-creating effect of currency unions is of considerable importance since it opens a new way for research. If a common currency does not only affect trade through a reduction in transaction costs, what would explain this effect? Until now, only a very few studies have searched for another way of modelling the link between trade and monetary union. Hence, a first step would consists in checking for the robustness of Rose's estimate, before trying to model the effect of a currency union on a stronger theoretical basis.

III. What is behind?

Following Rose's paper, several authors tried to check for the robustness of Rose's original finding. They generally choose either to apply the same method on a smaller data set, or to introduce more variables in the initial gravity equation. In this section, we will present the contributions which are directly connected to Rose's original article, *based on the same data set*[2]. All authors highlight some drawbacks in Rose's original estimation. For some, the major source of bias comes from the aggregation of heterogeneous kinds of currency unions. For others, it relies on a non-random selection of the currency unions included in the original sample. Finally, a third possible source of bias can be generated through the existence of omitted variables.

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[2] Several papers have also tried to estimate the effect of a currency union on trade, using other samples. Among these, Flandreau and Maurel (2001) focus on different monetary arrangements in 19th Century Europe, Thom et Walsh (2002) study the break-up of the monetary union between Ireland and the UK in 1979, and Micco, Stein and Ordoñez (2002) exploit very recent data on EMU countries. A lot of these references can be found in Rose (2002).

1. The aggregation bias

Several authors emphasized that an aggregation bias could potentially affect the original estimate. As highlighted by Nitsch (2002), Rose's sample covers multilateral currency unions as well as unilateral currency unions (for instance Panama dollarized since 1904) as well as little dependencies which have adopted the currency of their former colonial power (like Guadeloupe, or St. Helena)[3]. If these three groups of monetary unions affect trade differently, then the aggregation in a unique variable (the currency union dummy) will provide biased results.

One way to prevent this aggregation bias is to break up the original sample, i.e. to apply the same method to smaller homogenous data sets. Nitsch (2002) reduces the initial sample to two multilateral monetary unions: the CFA franc zone in West and Central Africa and the Eastern Caribbean Currency Union (consisting of eight small islands which were all British colonies). He justifies his choice arguing that these two multilateral monetary unions are homogenous in terms of production structures, historical experiences and cultural customs. The OLS estimation on this reduced data set provides new results: countries within the CFA zone would only trade 55% more with each other, while in the case of Caribbean countries, the estimate is still smaller and not significantly different from zero.

But one can also focus on unilateral currency unions. Klein (2002) studies the effect of a dollarization on trade. Hence, he eliminates all the observations, in Glick and Rose's data set, which do not include the United States[4]. He shows that the United States would trade 65% more with dollarized countries than with a typical non-union country. This estimated effect is far beyond the estimate of Glick and Rose (2002) or Rose (2000), and is significant only at the 90 percent level of confidence. He also studies the opposite case of countries which have adopted the dollar, and shows that these countries do not trade more with the United States than with other countries. Furthermore, the introduction of a binary variable indicating fixed exchange rates, increases the value of the monetary union variable, but without any

[3] Note that a multilateral union is characterized by the decision of several countries to create and share a *new* currency (e.g. EMU), while a unilateral union is defined by the decision of one or several countries to adopt the currency of another country (usually the dollar).

[4] Glick and Rose (2002) provide a larger data set including 217 countries from 1948 to 1997. Klein restricts this data set to the post Bretton Woods era. On this period, dollarized countries, whose data are complete, include Bahamas, Bermuda, Dominican Republic, Guatemala, Liberia and Panama.

significant difference between the two coefficients. Klein concludes to an identical effect on trade whatever the monetary arrangement (a fixed exchange rate regime or a currency union).

Another method to deal with this aggregation bias is to introduce in the initial equation different dummies to isolate each group of monetary unions. Levy Yeyati (2001) estimates the original gravity equation on the complete data set, and adds two dummies to distinguish between multilateral and unilateral currency unions[5]. He obtains a comparable coefficient for the case of unilateral currency unions (an increase in trade of 238%), but the estimated coefficient shrinks to 65% when considering multilateral currency unions (a result close to Nitsch's estimate). Levy Yeyati then concludes that the large coefficient obtained by Rose is due to the weight of unilateral currency unions in the original data set. As this group of currency unions consists mainly of "small subnational entities with strong institutional and political ties with the issuer of the common currency", the aggregate estimate has great chance to be over-estimated.

2. The self-selection bias

But the aggregation bias is not the only one which could affect the original estimate. The initial currency union groups can be non-randomly selected (Persson, 2001). Indeed, if there is a systematic correlation between the currency union dummy and other determinants of trade (captured by the explanatory variables), then the OLS estimation will provide biased results. In Rose's data set for example, countries belonging to a currency union are also generally smaller, poorer and geographically closer (Persson, 2001). Furthermore, Persson highlights a second potential source of bias. If the relation between some of the traditional determinants of trade – like trading costs - and the volume of trade is not linear, then the currency union dummy would capture the impact of these non-linear components, and therefore be over-estimated.

In order to prevent these two potential bias, Persson proposes an alternative strategy. This radically different method consists in creating a control group among countries which have their own currency, but which share the characteristics of countries belonging to a currency union. To create this control group, he constructs a "propensity score" evaluating the

[5] The first group includes the CFA franc zone and the Eastern Caribbean currency union, while the second group is defined as the unilateral adoption of the currency of another country.

probability of being put in the currency union's group, as a function of the explanatory variables. The comparison between country-pairs which share the same currency and country-pairs in the control group allows him to estimate the effect of a currency union on trade. His points estimates range from 13% to 65%.

In response to Persson, Rose (2001) argues that the methodology adopted for modelling currency union membership presents at least three weaknesses. First of all, Persson's strategy does not account for the time-series dimension of the data. Secondly, it lacks theoretical foundations, and finally, it fits and predicts poorly (for instance, it predicts that a Free Trade Agreement detracts from the likelihood of countries belonging to a currency union).

On a more restricted scale, one can also suspect the existence of a correlation between the currency union dummy and some political variables (Mélitz, 2001). To confirm this selection bias, Mélitz divides the Frankel and Rose's data set[6] into two groups. The first group consists in countries belonging to a monetary union *and* to a political union or a Free Trade Area (108 country-pairs), while the second group comprises the countries which *only* belong to a monetary union (56 country-pairs). The estimation of the gravity equation including these two binary variables – instead of the initial currency union dummy - gives similar results for these two extra-variables (the first coefficient is 2.18, while the second one is 1.87). Mélitz argues that if Rose's original estimate was reliable, then we would obtain a greater coefficient on the first variable, because this coefficient should reflect the impact of *both* political union or FTA and currency union. Hence, Mélitz's result suggests that the currency union dummy captures also the impact of political variables. However, unlike the previous authors, he does not reject all intrinsic effect, since, from his estimate, he calculates that a currency union could lead to a doubling in trade.

3. The omitted variable bias

The major problem could not be, as Persson argues, a problem of correlation between the currency union dummy and observable variables, but a possible correlation with *un*observable variables. If for example, cultural or historical ties between countries increase their propensity to form a currency union as well as their bilateral volume of trade, then the OLS estimate will

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[6] This sample is very close to Rose's original sample. It is slightly larger, since it integrates an additional year (1995).

be biased. To correct for this omitted variable bias, Tenreyro (2001) first estimates a selection model for currency union, based on the theoretical models of Mundell (1961) and Alesina and Barro (2002), to estimate afterwards simultaneously this model and a traditional gravity equation. Among the possible determinants of currency union membership, she considers comovement of output shocks, comovement in prices, the size of each country, the importance of trading costs, different measures of cultural similarity and a propensity to adopt inflationary policies[7].

Besides this econometrical aspect, Tenreyro tries to avoid a possible sample selection bias, coming from the fact that estimating a gravity equation in logarithm leads to eliminate automatically country-pairs with zero trade. As the pattern of trade between currency union members is often irregular – alternating years of large trade volume and years of very small trade – eliminating this kind of observations may lead to inflate artificially the estimated currency union coefficient. To address this sample selection problem, she averages trade flows over five-year intervals. After taking into account these two potential bias, the estimated effect of a currency union on trade shrinks to about 60%.

To address the same problem of unobservable variables, Rose and Van Wincoop (2001) apply another method which consists in introducing dummies for each country – whereas it is an exporter or an importer. These new binary variables should account for all factors explaining why a country has a particular propensity to trade - that is not yet explained by traditional variables like size or distance. This new estimate results in a slightly smaller effect of currency unions (a trade-creating effect of 136%).

However, this specification does not cover the general case, since it supposes that we take into account each country characteristics but *not* possible interactions between these two variables. Egger et Pfaffermayr (2000) argue that the omission of the interaction between exporter and importer dimensions, i.e. the omission of the bilateral fixed effect – a dummy which is 1 for the particular trade flow between i and j – could lead to biased estimates. To avoid all kind of omitted variable bias, these authors state that the estimated equation should take the following form:

(2) $Y_{ijt}=X_{ijt}b+q_i+q_j+l_t+a_{ij}+h_{ijt}$ where ij is a couple of countries, t the time index.

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[7] This last factor is proxied by a trade agreement variable. The reason for this choice is that the author supposes that countries which favour free trade policies will also be prone to more discipline, especially in terms of inflation.

q_i and q_j are dummies for each country, and a_{ij} is the bilateral fixed effect. The estimation of this equation is then equivalent to a specification where we just keep the fixed bilateral effect (a_{ij}) and the fixed time effect (l_t) (see Christensen, 1987).

Even if these two approaches succeed in highlighting an important source of bias in Rose's original estimation, they often fail to offer an alternative reliable estimation. Tenreyro describes *partially* currency union membership, since her model explains only 60% of the total variance. Hence variables not yet identified could continue to impose a bias on the estimation. The model proposed by Rose and Van Wincoop seems more appropriate since it does not require a precise identification of these unobserved characteristics, but as we have seen previously, this specification is only a particular case of a more general form.

In the next section, we will show that, even if it is impossible to identify these unobserved characteristics, it exists a method which eliminates all bias from the estimation.

IV. Towards a more reliable method of estimation: panel-data techniques

Until now, almost all estimations have been conducted using cross-sectional analysis[8]. Nevertheless, an estimation using panel-data techniques has two major benefits. First of all, it can account for the time-series variations as well as for the cross-sectional dimension of the data. Regarding the small number of currency unions, this additional dimension is not negligible. Moreover, it can control for unobservable characteristics which may be correlated with other observable variables. We argued in the previous section that these unobservable characteristics probably exist – they can take the form of political, historical or cultural ties between countries – and can be correlated to some of the explanatory variables (currency union membership for example).

1. The fixed effect model

Among panel-data methods of estimation, the fixed effect model has a major benefit: it always provides consistent estimates, even when there is a correlation between the

[8] With the exception of Pakko and Wall (2001) and Glick and Rose (2002) who proceed to within-groups estimations (see below).

explanatory variables and the specific effect. This model consists in estimating first differences, and consequently eliminates from the equation all variables that do not vary across time (including the bilateral specific effect). If the equation we want to estimate takes the following form[9]:

(3) $Y_{ijt}=X_{ijt}b+Z_{ij}g+a_{ij}+h_{ijt}$ where ij is a couple of countries and t the time index

Then the *Within* estimator removes all possibility of correlation between the explanatory variables (X et Z) and the bilateral specific effect (a_{ij}) since it leads to estimate:

(4) $Y_{ijt}-Y_{ij.}=\left(X_{ijt}-X_{ij.}\right)b+\left(h_{ijt}-h_{ij.}\right)$ where $Y_{ij.}$ $X_{ij.}$ et $h_{ij.}$ are the means of Y_{ijt} X_{ijt} and h_{ijt} on the period of observation.

To our knowledge, only two papers have applied this method to our purpose. Pakko and Wall (2001) use Rose's data set, and obtain a negative and non-significant coefficient for the currency union dummy. Glick and Rose (2002) use the same methodology applied to a different sample, and find a significant positive coefficient. Even if it is smaller than the initial coefficient, it indicates a still great effect since, according to this result, countries which enter in a currency union would see their bilateral trade increased by 90% (= $e^{0,65}$ – 1).

2. The random effect model

Contrary to the fixed effect model, the random effect model (GLS estimator) account for the time-series as well as for the cross-sectional dimension of the data. Therefore it can be more efficient, especially when we have little time-series variation. Moreover, it allows to estimate time-invariant variables – thanks to the cross-sectional dimension – but also to estimate with more reliability variables, as our currency union dummy, which do not vary much across time. But the counterpart of these advantages is that it could provide biased results if the specific effect is correlated to some of the explanatory variables.

Fortunately, the presence of this bias can be verified – using a Hausman test – and can even be corrected when it exists, thanks to the method proposed by Hausman and Taylor (1981). As we will see in the next section, this procedure derives consistent and efficient estimates even for time-invariant variables.

[9] To simplify the presentation, we will omit the fixed time effect in this section.

3. The Hausman Taylor method

Consider equation (3). The first step of the Hausman Taylor's procedure is to test whether it exists a correlation between unobserved characteristics – captured by the specific effect a_{ij} – and some of the explanatory variables. This can be done easily through a test that compares the random effect model to the fixed effect model (Hausman, 1978). If this correlation does not exist, then both model provide consistent estimators, but if some of the explanatory variables are correlated, then the only "right" model is the fixed effect model.

Hence, we test the null hypothesis H_0: $E(a_{ij} / X_{ijt}, Z_{ij}) = 0$ against the alternative hypothesis H_1: $E(a_{ij} / X_{ijt}, Z_{ij}) \neq 0$. If H_0 is rejected, we have to reformulate the cross-section specification to obtain a property of orthogonality.

The second step then consists in defining some instruments for the variables correlated to the specific effect. Hausman and Taylor divide the set of explanatory variables into four groups, depending on their possible correlation with the specific effect: X1 (containing k1 columns), X2 (k2 columns), Z1 (g1 columns) and Z2 (g2 columns) (see table 2).

TABLE 2 : CLASSIFICATION OF THE EXPLANATORY VARIABLES

	Not correlated to the specific effect	Correlated to the specific effect
Vary across time	X1 (k1)	X2 (k2)
Do not vary across time	Z1 (g1)	Z2 (g2)

The variables which need an instrumentation are X2 and Z2. Therefore, they define four instruments: deviations from group means of the variables X1, deviations from group means of the variables X2, means of the variables X1 and finally the time invariant variables Z1[10]. The first two instruments serve for the correlated variables X2, and the means of X1 serve as instruments for the time-invariant correlated variables Z2. It follows that a necessary condition for the identification of b and g is that the number of variables uncorrelated X1 should be at least equal to the number of variables Z2 (with our notations: k1≥g2). Note that an important contribution of this method is to derive instruments directly extracted from the

[10] The matrix of instruments takes the following form: [QX1, QX2, PX1, Z1], where Q is the matrix of projection transforming a vector X_{ijt} into a vector of deviations from group means ($X_{ijt} - X_{ij.}$) and P transforms a vector X_{ijt} into a vector of group means ($X_{ij.}$).

set of the explanatory variables. Since it may be difficult to find appropriate instruments, excluded from the regression, which are uncorrelated with the specific effect, this kind of instrumentation could be very useful.

This second step provides consistent estimators. But, because this procedure ignores the autocorrelation structure of the disturbance, these estimators are inefficient (Guillotin and Sevestre, 1994). Therefore, Hausman and Taylor propose to improve the model's efficiency by transforming the initial equation. Consequently, the third step of this procedure is to use the pre-defined instruments to estimate the following equation:

(5) $$Y_{ijt} - (1-q)Y_{ij.} = \left(X_{ijt} - (1-q)X_{ij.}\right)b + q\, Z_{ij}\, g + q\, a_{ij} + \left(h_{ijt} - (1-q)h_{ij.}\right)$$

Where $q = \sqrt{\dfrac{s_h^2}{s_h^2 + T s_a^2}}$ T is the number of periods of observation for each country-pairs, s_h^2 et s_a^2 are the variance components of the time-varying error component and of the specific effect, which can be easily estimated thanks to the *Within* and the *Between* models[11] (Guillotin and Sevestre, 1994). The instrumented estimation of this last equation provides consistent and efficient estimators, and when the model is over-identified (k1>g2), these estimators are even more efficient than the *Within* estimators. Hausman and Taylor summarize it, stating that: "the method is a two-fold improvement over the within-groups estimator: it is more efficient and it produces estimates of the coefficients of time-invariant variable" (p. 1393).

Finally, the last step consists in testing our assumption about the non-correlation of the variables X1. Hausman and Taylor propose to compare the efficient estimators provided by this method to the *Within* estimates. Under the null hypothesis of no-correlation between the instruments and the specific effect, both models provide consistent estimates. This test is then equivalent to a Chi-2 test with (k1-g2) degrees of freedom. Note that it can be run only if the model is over-identified (k1>g2).

In the case of an unbalanced data set, this method should be modified because the number of periods of observation (T) differs across country-pairs. Instead of (5), we shall estimate the following equation:

(6) $$Y_{ijt} - (1-q_{ij})Y_{ij.} = \left(X_{ijt} - (1-q_{ij})X_{ij.}\right)b + q_{ij}\, Z_{ij}\, g + q_{ij}\, a_{ij} + \left(h_{ijt} - (1-q_{ij})h_{ij.}\right)$$

[11] The *Between* estimator is given by the estimation on group means : $Y_{ij.} = X_{ij.}\, b + Z_{ij} g + a_{ij} + h_{ij.}$

Where $\theta_{ij} = \sqrt{\dfrac{\sigma_\eta^2}{\sigma_\eta^2 + T_{ij}\sigma_\alpha^2}}$

In this case, σ_η^2 and σ_α^2 cannot be estimated easily because the error components of the *Within* and *Between* models are heteroskedastic[12]. To estimate properly these two variances, Guillotin and Sevestre (1994) propose to reassemble country-pairs by groups in which the number of periods of observation is constant[13]. The rest of the procedure is not modified.

V. Our results

1. The omitted variable bias

We stated in the second section of this paper that a major source of bias lay in the existence of omitted variables. Persson (2001) suspects another kind of bias – related to the observable characteristics. But he also admits that a bias on the *un*observables could affect the estimation since he states on page 6: “If such omitted variables of the trade relation are also correlated with the common-currency indicator, then I certainly have a problem with selection on unobservables, but so has Rose”. Rose (2000), for his part, tries to minimize this kind of bias by integrating additional controls in his initial equation. But this method, as the one pursued by Klein or Nitsch, is too uncertain to be totally reliable. As noticed by Pakko and Wall (2001): “One cannot create variables to control for every pair-specific factor in the universe of trading pairs. The only sensible solution is to include a dummy variable for each pair that indicates trade between all pairs of trading partners, i.e., to create trading pair-specific fixed effects” (p. 40).

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[12] The variance of the error term of the *Within* model becomes: $V(\eta_{ijt} - \eta_{ij.}) = (T_{ij} - 1)/T_{ij}\,\sigma_\eta^2$, and for the *Between* model it becomes: $V(\alpha_{ij} + \eta_{ij.}) = \sigma_\alpha^2 + (1/T_{ij})\sigma_\eta^2$

[13] For each group k, one has to estimate the instrumented *Between* model from which we derive an estimation of the variance $V(\alpha_{ij} + \eta_{ij.})(k)$ and then apply OLS to the model: $V(\alpha_{ij} + \eta_{ij.})(k) = \sigma_\alpha^2 + (1/T_{ij})\sigma_\eta^2$, which provide, *in fine*, consistent estimators of σ_α^2 et de σ_η^2.

Indeed, the previous section showed that the only reasonable way to introduce these unobservable characteristics was to use panel-data techniques. This is why Pakko and Wall (2001) on the one hand, and Glick and Rose (2002) on the other hand, use the *Within* estimator to measure the impact of joining or leaving a currency union (see IV.1). This method derives consistent estimations but, as it ignores the cross-sectional nature of the data, it is not fully efficient - especially in the case of the currency union coefficient which do not vary much across time. As indicated previously, there are few cases of existing currency unions, but there are still less cases of creation or dissolution of monetary unions. Pakko and Wall's estimate of the currency union coefficient is based on 8 observations, while Glick and Rose's estimate relies on 146 cases of creation/dissolution. But even this last number is very small in comparison with the 219 558 observations of the entire sample.

Hence, the only feasible way to restore fully efficient estimators, without introducing a bias coming from the potential correlation of the specific effect with some of the explanatory variables, is to implement Hausman and Taylor method.

2. The estimation's results

We estimate the following model:

(7) $$\ln(X_{ijt}) = b_0 + b_1 \ln(Y_{it}) + b_2 \ln(Y_{jt}) + b_3 \ln(pop_{it}) + b_4 \ln(pop_{jt}) + b_5 \ln D_{ij} + b_6 \mathrm{Re}gional_{ijt} + b_7 Border_{ij}$$
$$+ b_8 Comlang_{ij} + b_9 Comcol_{ij} + b_{10} Comctry_{ijt} + b_{11} Colonial_{ij} + b_{12} V(e_{ijt}) + b_{13} CU_{ij} + l_t + a_{ij} + e_{ijt}$$

where l_t is the time fixed effect, a_{ij} is the bilateral specific effect and e_{ijt} is the error time-varying component. We will suppose, as usually, that this component is uncorrelated with the explanatory variables, and that it has zero mean and constant variance. For the meaning of the other variables, see section II. Rose's trade data come from the United Nations Statistical Office. Populations and real GDP per capita are taken from *Penn World Table 5.6* and World Development Indicators (World Bank). He calculated great-circle distance using latitude and longitude taken from the CIA's *World Factbook*. Other dummies come from the same source. The dummy of regional trade agreements is constructed using WTO's web site. The variable of volatility is measured as the standard deviation of the first difference of the monthly logarithm of the bilateral nominal exchange rate in the five years preceding the current period (calculated using IMF data).

In our specification, we choose to separate the GDP and the population of the two countries – instead of entering the product of both variables – for two reasons. Firstly, it is a more conventional and a more general form of the gravity equation, and secondly, it provides more variables which can be potentially used as instruments. We regress this equation on the initial Rose's data set, previously transformed to be balanced – i.e. reduced to the country-pairs whose data are available in the five years of the sample. The results of the OLS estimation are given in the first column of table 3. The variables of population can have in theory positive or negative effects on trade. An important population can reflect a larger market, and therefore can enhance trade. Conversely, it can also reveal a relative self-sufficiency. However, in several empiric studies, this negative effect seems to dominate (Carrere, 2001). In our estimation, both variables are negative and insignificant. Except for this ambiguous variable, the gravity model works well: all coefficients have the expected sign, and are largely significant. For example, more variability in exchange rates reduce bilateral trade, whereas a Free Trade Agreement enhances trade between member-countries. These results are generally close to Rose (2000), even if the currency union dummy is slightly larger.

First of all, we need to test if there is a correlation between the specific effect ($á_{ij}$) and some of our explanatory variables. Therefore, we estimate the fixed effect and the random effect model (GLS), and compare the results given by these two estimations. Table 3 shows that some of our variables are indeed endogenous[14], since the Hausman statistic is large and significant. These two models give sensibly different estimates, especially in the case of the population variables and of the Free Trade Agreement dummy.

Consequently, Hausman and Taylor method should be implemented. We have made two different assumptions about the status of our explanatory variables. In a first step (hypothesis a), we suppose that the Free Trade Agreement dummy and the currency union variable are exogenous. The intuition behind this assumption is that the decision to reach a trade agreement or to join a currency union[15] is strictly political, and hence exterior to our model. This intuition has been followed by several authors. In particular, Flandreau and Maurel (2001) state on page 12: "We argue that monetary coordination is largely a political decision,

[14] We will define here an endogenous variable as a variable which is correlated to the bilateral specific effect.

[15] Which can be seen as a deeper stage in regional integration.

whose motivation it is hard to ascribe to strictly economic motives, and which is reversed much less often than economics would warrant. In other words, by contrast with the traditional OCA approach, we treat monetary unification as a mostly *exogenous* decision". This assumption allows them to derive a small model in which monetary coordination favors business cycle correlation, which in turn leads to a reduced current account constraint, and finally to more trade integration. Instead of considering both the Optimum Currency Area criteria and the monetary union creation as endogenous, they prefer to presume that the only causality goes from monetary union to OCA criteria.

Under this hypothesis, our endogenous variables are limited to the GDP, the population and the volatility of exchange rates. The effect of currency union on trade seems confirmed, and the estimate is even more important. The Free Trade Agreement dummy is also positive and significant at 3%, even if its value is considerably smaller compared to the OLS or GLS estimates. A part of the correlation between our explanatory variables and the specific effect had been removed, since the Chi-2 statistic is lower (12,43 compared to the initial 109,89). However, this statistic is still significant at 10%, suggesting that the model is not yet well specified. This requires the modification of our initial hypothesis.

Hence, in a second step (hypothesis b), we try to restore the endogeneity of these two variables. Indeed, a lot of authors have argued that the reaching of a trade agreement or the creation of a currency union can be explained – at least partially – by institutional, cultural or historical ties. It is the reason why Tenreyro (2001), for instance, constructs a model explaining currency membership before any estimation of the effect of a common currency on trade. This same argument leads Pakko and Wall (2001), and Glick and Rose (2002) to use a *Within* estimator. To put it plainly, let advance an example. With this assumption we would presume, that the use of the US dollar in Panama and the particular pattern of bilateral trade between the US and Panama relies on historical grounds. This seems plausible.

The results are accordingly reversed, since the regional integration and the currency union variables loose their positive and significant coefficients. We have computed two different estimates. In the first one (column 5), we assume that the GDP and the population variables are exogenous. This allows us to reduce the number of endogenous variables to three (the trade agreement, the currency union dummy, and the volatility of exchange rates variable). In the second estimate (last column), by contrast, we try to maximize the number of endogenous variables, and keep only two exogenous variables[16] – the minimum that allows us to test our model (see IV.3.). This second estimation gives much better results since the Chi-2 statistic is now very low (= 3,97) and statistically insignificant, indicating that we succeed in eliminating all correlation with the specific effect. Note that all coefficients are now very close to the *Within* estimate. Quantitative result regarding the currency union coefficient should not be interpreted too literally since the standard error is so high that, for example in the second estimation, the interval describing its value at the 95 percent level of confidence is: [-896,67 ; 823,53].

However, in both cases, the important result is the same: membership in a regional integration regime cannot, *per se*, affect trade. This result could surprise when referring to the impressive work of specification conducted by Rose. However, it is consistent with other contributions (see III.1 and III.2).

Our results suggest that Rose's positive and significant coefficient is essentially due to a bias coming from omitted variables that are correlated with the currency union dummy. When we do *not* control for this correlation (hypothesis a), the currency union effect still exists and is even larger. But after correction (hypothesis b), it appears that the currency union variable is neither positive nor significant. The different contributions presented in section III, appear in between these two estimates, since they generally find a smaller and less significant coefficient, but proceed only to a partial correction of the existing bias.

Another important result can be derived from this second estimation. Once we introduce a bilateral specific effect and control for a possible correlation, the estimated coefficients on the distance, common land border, common language and common history variables become statistically insignificant. This could be explained by the fact that in the OLS estimates, these variables were capturing a part of the unobservable characteristics, now included in the specific effect. This result can be found in other studies which apply Hausman and Taylor procedure (see for example De Sousa and Disdier, 2002). Therefore, another benefit of this method is to address the possible problem of misspecification of these variables. Especially, in the case of the distance, it is often argued that this variable is a bad proxy for trading costs (see Pakko and Wall, 2001 for further discussion).

[16] The population of country j and the dummy of common country.

Finally, when compared to the OLS estimation, the estimated coefficient on exchange rate volatility shrinks to one third.

TABLE 3: HAUSMAN AND TAYLOR METHOD OF ESTIMATION

	OLS	WITHIN (fixed effect model)	GLS (random effect model)	HT1 (1) (hypothesis a)	HT2 (2) (hypothesis b)	HT2' (3) (hypothesis b)
Currency Union	1.89*** (.28)	-	1.93*** (.51)	3.01*** (.96)	-22.12 (19.81)	-36.57 (438.79)
Volatility of exchange rate	-.015*** (.002)	-.005*** (.0019)	-.005*** (.0017)	-.0047*** (.0018)	-.0034** (.0017)	-.005*** (.0019)
Gdp_i	1.33*** (.016)	1.28*** (.06)	1.32*** (.03)	1.28*** (.06)	1.31*** (.04)	1.24*** (.06)
Gdp_j	1.64*** (.018)	1.33*** (.07)	1.59*** (.03)	1.35*** (.06)	1.44*** (.06)	1.34*** (.07)
Pop_i	-.56 (.019)	-.19 (.13)	-.56*** (.03)	-.36*** (.11)	-.55*** (.05)	-.43*** (.06)
Pop_j	-.87 (.019)	.14 (.14)	-.82*** (.03)	-.02 (.13)	-.70 (.05)	.15 (.15)
Distance	-1.03*** (.02)	-	-1.06*** (.04)	-1.29*** (.08)	-1.23*** (.14)	-1.63 (3.10)
Regional trade agreement	.84*** (.09)	.11 (.11)	.36*** (.01)	.22** (.10)	.12 (.10)	.10 (.11)
Common language	.44*** (.04)	-	.43*** (.08)	.74*** (.15)	.63*** (.21)	1.15 (4.15)
Common border	.30*** (.10)	-	.29 (.19)	-.45 (.37)	-.19 (.49)	-1.22 (8.53)
Common colonizer	.88*** (.06)	-	.85*** (.12)	.95*** (.24)	1.03*** (.29)	1.34 (5.39)
Common country	.44 (.73)	.45 (.58)	.52 (.56)	.49 (.54)	.55 (.55)	.46 (.59)
Colonial relationship	1.84*** (.11)	-	1.87*** (.20)	1.17*** (.42)	1.79*** (.37)	.56 (3.58)
N° of obs.	12 930	12 925	12 925	12 925	12 925	12 925
R^2	0.68	0.45	0.68	0.23	0.22	0.10
Theta				0.18	0.18	0.18
Hausman Test			109.89*** Chi-2 (11)	12.43* Chi-2 (6)	47.04*** Chi-2 (8)	3.97 Chi-2 (5)

Notes:
Constant and year-specific fixed effects not reported. Standard errors in parentheses.
* significant at 10%, **significant at 5%, *** significant at 1%.
(1) Endogenous variables: Gdpi, Gdpj, Popi, Popj, volatility (X2).
hypothesis a): cu and regional trade agreement exogenous.
(2) Endogenous variables: volatility, regional (X2) ; cu (Z2)
hypothesis b): cu and regional trade agreement endogenous.
(3) Endogenous variables: volatility, regional trade agreement, Gdpi, Gdpj, Popi (X2) ; cu (Z2)

Hence, currency union membership would not have any effect, *per se*, on the volume of trade. This does not mean obviously that it has no effect at all. On the contrary, we think that the results we obtained reveal that, for our concern, the gravitational approach is maybe not the right one. There should be *some* effects on trade, but these effect are probably more indirect, coming from an impact on interest rate, capital flows, or, as argued by Flandreau and Maurel, the current account constraint. The gravity equation is therefore too restrictive to account for all these aspects.

VI. Conclusion and suggestions for future research

We agree with Rose when he says that "even if we don't know why a common currency makes a difference, it is plausible that it does" (p. 23), but nevertheless we do not think that the effect he derived is the *real* effect. On the contrary, we argued that the gravitational model is not able to reveal a strong effect, once controlling for unobservable characteristics. Consequently, we think that a gravity equation is maybe not a sufficiently complete framework for modelling this kind of monetary arrangements, and that we should use instead a much more general model, based on stronger theoretical foundations.

To construct this model, one has to understand better what is behind the creation of a currency union. The effect of a currency union on trade may not come from a reduction in transaction costs – as stated by the OCA approach -, but could be connected to other factors, financial integration for example. This issue has not been totally capitalized, since until now, the major part of the economists has followed the steps of Mundell in modelling currency unions. Moreover, it is of major importance since many countries in the world and in Eastern Europe in particular, are contemplating adopting a common currency or the currency of other countries.

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