Économie internationale 115 (2008), p. 193-212

ÉCONOMIE
INTERNATIONALE
INTERNATIONAL ECONOMICS

Currency transaction tax elasticity: an econometric estimation

Francis Bismans & Olivier Damette[1]

Article received on April 30, 2008
Accepted on October 16, 2008

Abstract. This article endeavours to measure the elasticity of the volume of the currency exchange transactions to a tax on them. The analysis is principally based on cointegration techniques. This paper is the first attempt to estimate the influence of a currency transaction tax on the foreign exchange market volume trading. The econometric estimations suggest that the *forex* trading volume could be significantly reduced by a Tobin tax. Nevertheless, elasticities are heterogeneous with respect to the currency pairs: the largest elasticities are the euro/dollar and sterling/dollar currency pairs that are also the most traded exchange parities. The values of the estimates are lower when the SURE (Seemingly Unrelated Regression Equations) estimator is used than when the panel estimation is implemented.

JEL Classification: C22; C23; F31; H20.
Keywords: Foreign Exchange Market; Currency Transaction Tax; Elasticity; Cointegration; SURE; Panel Data.

Résumé. Cet article cherche à mesurer l'élasticité du volume des transactions de change à un impôt qui leur serait affecté. L'analyse repose essentiellement sur les techniques de cointégration. Il s'agit de la première contribution visant à estimer l'effet qu'aurait un tel impôt sur le volume des transactions sur les marchés des changes. Les estimations économétriques montrent que les volumes échangés sur le marché des changes pourraient diminuer de manière significative en cas d'application d'une taxe Tobin. Toutefois, les élasticités varient selon les devises pour lesquelles elles sont calculées : les plus élevées sont obtenues pour les couples euro/dollar et sterling/dollar, qui sont les parités les plus échangées. Les résultats des estimations sont plus faibles si l'on recourt à l'estimateur SURE (Seemingly Unrelated Regression Equations) plutôt qu'à un calcul sur données de panel.

Classification *JEL* : C22 ; C23 ; F31 ; H20.
Mots-clefs : Marché des changes ; taxe sur les transactions de change ; élasticité ; cointégration ; SURE ; données de panel.

1. Corresponding author: Olivier Damette, PhD in economics & ATER (Attaché temporaire d'enseignement et de recherche), Université Nancy2; Research fellow at BETA-Bureau d'Économie Théorique et Appliquée-, CNRS UMR 7522, (olivier.damette@univ-nancy2.fr).
Francis Bismans, PhD in economics, Professor in quantitative economics, Université de Nancy.

1. Introduction

In 1972, during his Janeway Lectures at Princeton University, Tobin first proposed to tax all spot transactions in the foreign exchange market. According to him, the tax would be assessed on each conversion of domestic currency into any foreign currency at a low level tax rate (1% in 1972). Although the well-known Tobin tax was again suggested by Tobin in 1974 and 1978, this proposition was not greeted with enthusiasm by the community of the economists until 1990's. Consequently, the so called Tobin tax has experienced a chaotic existence. Due to the multiplication of financial crises (Mexico, East Asian, Argentina, Russia…) and of its solidarity relevance, the Tax became indeed surprisingly popular.

Nowadays, lots of new global taxes have been proposed (Carbon and environmental taxes, airline tickets tax…), but the Tobin tax is still in minds. For its supporters, this measure presents several advantages, such as increasing national monetary autonomy, reducing noisily short term flows leading to exchange rate instability, and raising revenues for multilateral projects. However, those effects have been more often ideologically analyzed and there are only few academic studies, despite the major and seminal contribution of Haq Kaul and Grunberg (1996). The theoretical and analytical studies are reduced to the models of Frankel (1996) whose background is the noise trading approach,[2] Bird and Rajan (2001) who extended the Frankel's framework, Bosco & Santoro (2004) and Eichner & Wagener (2005) who suggested a mean variance approach of the Tobin tax effects on the *Forex*.

At the same time, a lot of reports have been carried out by ministries (French, 2000; Belgian, 2001; Finnish, 2001; German, 2002) and institutional (European Parliament, 2000; OECD; 2002). A lot of studies produced various estimates of potential revenues from a Tobin tax (the most discussed question about this topic, see Felix and Sau, 1996; Spahn, 2002; or Nissanke, 2003). Nevertheless, they suffer from numerous drawbacks: notably the impact of a Tobin tax on the volume trading (that is the tax base) is not precisely taken into account.

All in all, there is finally a lack of academic and notably empirical studies about the Tobin tax as bought to the fore by the OECD (2002, p.194) because there is *"a lack of empirical evidence on the sensitivity of trading volumes with respect to the spreads"*, by Aliber et *al.* (2003, p.1) for whom *"the empirical evidence on the effect of transactions tax volatility price is rare"* and by Eichner and Wagener (2005, p.11) for whom *"theoretical considerations provide little guidance (…) the true effects of a Tobin tax can only be assessed empirically"*.

The aim of this paper is to propose an econometric framework in order to estimate those effects and to fulfil the lack of empirical studies about the Tobin tax topic. Indeed, it would be useful to quantify how the *forex* trading volume would decrease due to a currency transaction tax. This estimation is also very interesting in order to investigate the effect of such a tax

2. See Frankel and Froot (1990), De Long *et al.* (1990) and Jeanne and Rose (2002) for a noise trading approach from the foreign exchange market.

on the exchange rate volatility.[3] More precisely, our aim is to combine non stationary time series and panel data cointegration procedures in order to estimate the so called "Tobin tax elasticity".

The article is organised as follows: section 2 describes the data set and some methodological issues; section 3 exposes the results of time series cointegration; section 4 deals with the SURE estimates of the data; section 5 looks at panel data estimations (using the most recent panel unit root tests with cross section dependence) and section 6 concludes.

2. Data and methodology

Given that the currency transaction tax does not exist, we have to use a *proxy* to take it into account. In accordance with the *traders*, we assume that the Tobin tax implementation is equivalent to a rise in the transactions costs i.e. that currency transaction tax has the same effect as an increase in transactions costs. The Tobin tax literature has reached a consensus on this issue. In the foreign exchange market, transactions costs are reflected in observed *bid-ask spreads*. The *spreads* are the source of profits for *Forex* intermediaries and are currently at a very low level (see Spahn, 2002) for currently observed *spreads* reports). Even a Tobin tax of 0.05% is above the highest realized spreads in the Euro/dollar interbank market.

Therefore, in order to estimate the impact of a currency transaction tax on the trading volume, we estimate the relationship between the *volume* (v) and the *bid-ask spread* (s). Our baseline model is the following one:

$$\ln(v)_t = \text{constant} + E\ln(s)_t + \beta\ln(\text{other variables})_t + \text{error}_t,\ t = 1, \ldots, T \qquad (1)$$

where T is the time dimension and E denotes the elasticity.

To run this estimation, we need data for all the variables quoted above. The most important feature of the foreign exchange market is that it is decentralized and opaque. Consequently, times series on actual trading volume and prices are very difficult to obtain (for instance Aliber and *al.*, 2003, used futures data as a proxy because those data are traded on a centralized market). Nevertheless, we use an original data set with the support of Reuters from *Reuters Dealing 3000*, which is an electronic platform used by interbank traders and the most popular foreign exchange market information providers with *Electronic Broking System* (EBS). Thus in this paper, we are working with real data of the foreign exchange market which is an improvement compared to the few existing databases (which used futures data or Reuters tick frequency as a proxy for trading volume).[4]

3. As explained by Haberer (2004), there are two effects on exchange rate volatility when imposing a Tobin tax. The tax reduces excessive volatility in highly speculative markets with high trading volume when the tax rate is small. In illiquid forex markets however, the tax might raise the volatility. Therefore, it is very important to evaluate the liquidity level of the market after introducing a Tobin tax.

4. See Hartmann (1998) for a survey considering the different existing databases about the foreign exchange market.

Our data set consists of intraday time series over the sample period from November, 24, 2004 to November, 25, 2004 for four currency pairs: dollar/euro (EUR/US$ hereafter), dollar/yen (JPY/US$ hereafter) dollar/sterling (GB£/US$ hereafter) and dollar/Canadian (CAD$/US$ hereafter).[5] The sample period shows the foreign exchange market working in a steady period (there was not minor or major crisis during the sample period, which ensures some stability in the data). Note that in terms of currency composition of *forex* turnover in 2004,[6] the dollar/euro (28%), dollar/yen (17%) and the dollar/sterling (14%) are the most traded currency pairs. Those three currency pairs account for near 60% of the total trading volume of the foreign exchange market. The dollar/Canadian is a minor currency pair, the sixth most traded (4%) and will be a mean of controlling the results. All in all, our data set is quite representative of the whole *forex* market.

For each currency pair, we obtain the following variables so transformed:

- the accumulative traded volume of payments in a currency pair, defined in billions of current US dollars (*volume*);
- the foreign exchange market volume for each minute (*volume2*);
- the quoted bid price, recorded for each minute (*bid*);
- the quoted ask price, recorded for each minute (*ask*);
- the spread between ask and bid (*spread*).

Given our series, our baseline model (1) can be rewritten as (2). Hence, the equilibrium Tobin tax elasticity coefficient for the three currency pairs can be obtained by the following log linear model estimation:

$$L volume_{it} = \beta_0 + \beta_1 \; Lask_{it} = \beta_2 \; Lspread_{it} + u_{it}, \; i = 1, \ldots N, \; t = 1, \ldots, T \quad (2)$$

where L denotes the base e-logarithmic function. Therefore β_2 can be directly interpreted as an elasticity coefficient and we expect it to be negative.

3. Time series analysis

Since our panel data set is very specific (N=4, T=1493), we test first for the existence of a cointegration relationship for each cross section, that is each currency pair, in the time dimension only. At first, we test for the non stationarity of the series with the usual time series unit root tests, i.e. Augmented Dickey Fuller (ADF, 1979), Dickey-Fuller Generalized Least Squares (DF-GLS, 1996), Kwiatkovski-Phillips-Schmidt-Shin (KPSS, 1992), and Elliot Rothenberg Stock Point-Optimal (ERS, 1996) tests.[7] Note that both Dickey Fuller tests are based on 30 lags, because we think that a large number of lags (i.e. only 30 minutes here) is

5. There is no particular reason to study this specific data and exclusively the four currency pairs mentioned above but these are the only data over the same time period we have after crossing the different currency pairs of our database.

6. Note that the last turnover by currency pair (published by the BIS in 2007) is near of 2004: 27%, 13%, 12% and 4% respectively.

7. As those usual tests are nowadays well-known, it is not relevant to develop them here.

necessary to take into account the presence of some high order autocorrelation in our series.[8] Nevertheless, in order to check the robustness of our unit root test results to this economic insight, we perform besides the KPSS test (where stationarity is the null hypothesis) and the ERS test (which is more powerful than the DF tests and has been found to dominate other unit root tests under certain conditions) by choosing the automatic lags selection with the Akaike criterion. All the results are reported in TABLE 1.

Table 1 - Results from unit root tests

Currency pairs	Variables	DFA	DF_GLS	KPSS	ERS
EUR	Lvol	−2.11 (0.23) ***	−1.07***	1.50***	7.46 (25)***
	Lspread	−2.75 (0.06)**	−1.78**	2.32***	8.88*** (27)
	Lask	−2.16 (0.21)***	0.42***	3.76***	40.94*** (5)
GB£	Lvol	−2.44 (0.13)***	−0.52***	0.66**	56.54*** (30)
	Lspread	−2.75 (0.06)**	−0.50***	1.70***	44.10*** (25)
	Lask	−2.75 (0.06)**	0.57***	3.00***	60.59*** (3)
CAD$	Lvol	−1.41 (0.15)***	−4.60	1.66***	5.87*** (60)
	Lspread	−0.05 (0.66)***	−6.21	1.51***	2.91** (18)
	Lask	−3.98 (0.02)	−0.80***	1.98***	12.81*** (7)
JPY	Lvol	−2.64 (1.00)***	2.37**	1.16***	6.64*** (30)
	Lspread	−12.51 (0.00)	−6.57	1.90***	6.75*** (29)
	Lask	−1.69 (0.43)***	−0.34***	1.83***	10.89*** (17)

Notes: Lag length is set to 30 in the DFA (with intercept) test and in the DF_GLS. No time trend is included in the model specification considering the charts of our series.
(.) are p-values with regards to the DFA and lags (according to Akaike criterion) with regards to ERS.
***, **, * indicate unit root at 1%, 5%, 10% significance respectively.

Overall, Dickey Fuller unit root statistics for the three variables imply not rejecting the null hypothesis of unit root in the EUR/US$ and GB£/US$ series at 1%. Each variable has a unit root and is thus integrated of order one I(1). It is however less obvious for the CAD$/US$ and JPY/US$ currency pairs (maybe the autocorrelation would be higher for the Canadian and Japanese currency pairs than for the others). For the CAD$/US$, ADF results indicate that all variables except *Lask* are level stationary, but DF-GLS outcomes show that only *Lask* has a unit root. The results are nonetheless clear-cut considering KPSS and ERS tests since all the variables have a unit root.

Since all the series are I (1), we now investigate the presence of one or more cointegration relationship(s). Indeed, since Granger (1983) and Engle and Granger (1987), it is well-known that if two or more series are integrated of the same order *d*, they are said to be cointegrated if it exists a linear combination between these series which is stationary. The

8. Selecting the lags according to the Akaike criterion leads to very close results to those obtained with 30 lags. This suggests an absence of specification error in the unit root tests in terms of selection of the lag length.

Engle and Granger procedure is limited to the investigation of the existence of only one chosen equilibrium relationship between our variables. Nevertheless, it is possible that a number of *k*-1 relationships exist between *k* variables, that is *volume*, *spread* and *ask* in our model.

Using the Johansen (1988) methodology, we test for the existence of the exact number of cointegrating relationships in a multivariate VAR (Vector AutoRegressive) model. We conclude that there exists one possible cointegrating relation for the EUR/US$ and the JPY/US$ currency pairs at 1% level but two cointegrating relations for the GB£/US$ and the CAD$/US$ currency pairs (Trace and Eigenvalue tests reproduced in TABLE 2a). For instance, the model (3) provides estimates of the cointegrating relationship relating to the EUR/US$ currency pair.[9] It is known that the cointegrating vector is not identified unless we impose some arbitrary normalization. The equation (3) reports estimates of the cointegrating vector relating to the EUR/US$ currency pair based on the normalization of the *Lvolume* variable. This equation shows that the normalized cointegrating coefficients have the expected sign (the *Lspread* coefficient is negative):

$$Lvolume2 + 11.48\ Lspread - 54.93\ Lask + 108.57 = 0 \qquad (3)$$

Consequently, the cointegrating relationship (3) is in line with the model (2). The Engle and Granger method is then performed to test the null hypothesis of no cointegration against the alternative of stationary residuals. In this way, our baseline model is the model (2) with time dimension only. TABLES 2 present the outcomes of the Engle and Granger methodology.

Table 2a - Results from Johansen cointegration tests

Hypotheses			EUR/US$		GBP/US$		CAD/US$		JPY/US$	
H0	H1*	H1**	λ_{trace}	λ_{max}	λ_{trace}	λ_{max}	λ_{trace}	λ_{max}	λ_{trace}	λ_{max}
$r = 0$	$r \geq 0$	$r = 1$	48.30	40.84	65.86	65.86	74.38	39.58	55.50	43.94
			(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
$r \leq 1$	$r \geq 2$	$r = 2$	7.28	6.22	20.72	20.72	34.40	29.39	11.57	11.01
			(0.29)	(0.31)	(0.00)	(0.00)	(0.00)	(0.00)	(0.07)	(0.05)
$r \leq 2$	$r \geq 3$	$r = 3$	1.06	1.06	1.77	1.77	5.40	4.97	0.55	0.55
			(0.35)	(0.35)	(0.21)	(0.21)	(0.48)	(0.51)	(0.52)	(0.52)

Notes: p-value in parentheses (.); * for Trace Test and ** for Max Test.
Lag length is set to 16 in EUR/US$, JPY/US$ and GBP/US$ tests, 3 in CAD/US$. Lag order is selected by AIC (Akaike), SC (Schwarz), HQ (Hannan-Quinn), FPE (Final Predictor Error) and LR (Likelihood Ratio) criteria.

9. To save place, we do not show here all the cointegrating relationships for all the currency pairs.

Table 2b - Estimation and residuals with Engle and Granger methodology

	Constant		Lask		Lspread		Cointegration test	
	Coeff.	p-value	Coeff.	p-value	Coeff.	p-value	Stat.	p-value
EUR/US$	126.44	0.000	−487.53	0.000	−0.61	0.000	−3.638	0.003
GBP/US$	−78.01	0.001	109.37	0.002	−0.55	0.000	−4.317	0.000
CAD/US$	−87.16	0.000	−482.17	0.000	−0.30	0.000	−5.578	0.000
JPY/US$	437.06	0.002	−94.815	0.001	−0.79	0.000	1.902	0.987

The last column of TABLE 2b (ADF) shows that all p-values are close to zero with the exclusion of the last line, i.e. the residuals are stationary for each currency pair except the JPY/US$ currency pair. For this reason, there is one cointegration relationship for each currency pair quoted above. Moreover, several results can be derived from the TABLE 2a:

- as expected, the coefficient value associated to the *Lspread* variable is negative for the four exchange parities: the logarithm of trading volume is decreasing in the logarithm of *spread*;
- the *Lask* coefficient in the sterling regression is surprisingly positive. This result may be explained by the coefficient value instability after implementing the CUSUM (Cumulative Sums of the Recursive Residuals) statistic;
- the coefficients for both EUR and GBP currency pairs are significant and very close to each other: –0.61 and –0.55. Therefore, an increase of 1% in transactions costs lead to a decrease of –0.61% and –0.55% of the Euro/dollar and of the Sterling/dollar volume trading respectively;
- lastly, as we could expect, the elasticity value for the Canadian dollar is lower than for the other currency pairs. We can explain this result by the low trading volume of this currency pair in the overall foreign exchange market trading volume. This type of transactions would be less sensitive to a one or two basis points (0.01% to 0.02%) currency transaction tax.

To check the robustness of those different results, we performed some other regressions to address if the results are sensitive to changes in the estimation. To this order, we conducted a linear regression version of the model (2) to check the coefficient signs (*Lspread* is always negative and significant at 1%) and we recomputed especially the logarithmic model (1) over two different subsets of the overall data set (we separated the sample to the 825^{th} value, i.e. when moving from first to second market day in the database). The values of the *Lspread* coefficients were slightly modified but were very close to each other over the two samples of time period (for instance, the *Lspread* values about the GBP/US$ were –0.55 and –0.53 respectively).

Finally, we assess the possibility of multicollinearity among *Lspread* and *Lask* by performing the Variance Inflation Factor (VIF) and the Tolerance tests. The square root of the VIF indicates how much larger the standard error is, compared to what it would be if the studied variable (*Lspread* or *Lask*) was uncorrelated with the other variable of the model (2). As outlined by the VIF values (1.04 for EUR/US$, 1.03 for GB£/US$, 1.00 and 1.00 for CAD$/US$ and JPY/US$ respectively) for *Lask* and *Lspread*, it is clear-cut that there is no mulitcollinearity among *Lspread* and *Lask*.

4. The SURE estimator

The estimations in TABLE 1 were based on the assumption that the error terms in the equations are independent. The intuition behind this hypothesis is that the behaviour of the traders is not the same across exchange rate currency pairs: the elasticities are therefore heterogeneous as outlined in the preceding section. However there is eventually a relation among exchange parities (denoted i and j), because traders react the same way to some news (macro and financial) for different currency pairs, so that the perturbations are likely to be correlated. Moreover all of the currency pairs are expressed in dollars.

This kind of correlation is precisely a contemporaneous one. Formally we have:

$$Cov\,(u_{it}\,,\ u_{jt}) = \sigma_{ij} \neq 0, \quad i \neq j$$

Adding such contemporaneous correlation into the model introduces additional information which is not incorporated when we estimate the four equations by OLS. Thus, it is possible to obtain a better estimator than the OLS estimator (section 3) by viewing equations as parts of a system. When the disturbances in a specific equation are contemporaneously correlated with the disturbances in other equations, such systems are known – see Zellner (1962) – as seemingly unrelated regression equations (SURE).

In our example, each equation may be written on the form:

$$\mathbf{y}_m = \mathbf{X}_m\,\boldsymbol{\beta}_m + \mathbf{u}_m\,, \quad m = 1\,,\ \ldots\,,\ 4\,,$$

where $\mathbf{y}_m$ and $\mathbf{u}_m$ are $T \times 1$ the matrices $\mathbf{X}_m$ of order $T \times 3$ and $\boldsymbol{\beta}_m$ is a 3×1 vector of parameters.

If $\mathbf{X} = \text{diag}\,(\mathbf{X}_1\,,\ \ldots\,,\ \mathbf{X}_4)$, $\mathbf{u} = (\mathbf{u}_1\,,\ \ldots\,,\ \mathbf{u}_4)'$ and $\boldsymbol{\beta} = (\boldsymbol{\beta}_1\,,\ \ldots\,,\ \boldsymbol{\beta}_4)'$ denote the system of four equations in stacked form as:

$$\mathbf{y} = \mathbf{X}\boldsymbol{\beta} + \mathbf{u}. \tag{4}$$

If the errors are contemporaneously correlated, the covariance matrix $\mathbf{V}$ in (4) is no longer diagonal. In fact, the correlations between the errors may be characterized by writing:

$$E\,(u_{it}\,u_{jt}) = \sigma_{ij}\,,\ i\,,\ j = 1,\ \ldots\,,\ 4, \quad i \neq j.$$

However, all non-contemporaneous disturbances are independent so that $E\,(u_{it}\,u_{js}) = 0$, for all $t \neq s$.

Consequently, the matrix $\mathbf{V}$ of order $4T \times 4T$ in (4) is given by:

$$\mathbf{V} = E(\mathbf{u}\mathbf{u}') = \boldsymbol{\Sigma} \otimes \mathbf{I}_T\,, \tag{5}$$

where $\boldsymbol{\Sigma} = (\sigma_{ij})$ is a 4×4 matrix and $\otimes$ denotes the tensor product.

Zellner's SURE estimator of $\boldsymbol{\beta}$ in (4) is got by feasible generalized least squares:

$$\begin{aligned}\hat{\boldsymbol{\beta}}_{SURE} &= (\mathbf{X}'\hat{\mathbf{V}}^{-1}\mathbf{X})^{-1}\mathbf{X}'\hat{\mathbf{V}}^{-1}\mathbf{y} \\ &= (\mathbf{X}'(\hat{\boldsymbol{\Sigma}}^{-1}\otimes \mathbf{I}_T)\mathbf{X})^{-1}\mathbf{X}'(\hat{\boldsymbol{\Sigma}}^{-1}\otimes \mathbf{I}_T)\mathbf{y},\end{aligned} \tag{6}$$

with $\hat{\boldsymbol{\Sigma}}$ a consistent estimator of $\boldsymbol{\Sigma}$.

TABLE 3 gives the values for Zellner's estimation procedure. All the coefficients are significant at 1% except for the *Lspread* variable in the JPY/US$ relationship. Moreover, the *Lspread* coefficients are generally of the correct negative sign. Comparison of the OLS cointegration results (section 2) and SURE results indicates that there is a general improvement in the efficiency of the estimates, evidenced by decreases in the p-values of all variables. The elasticities coefficients are smaller in the SURE equations. Unlike OLS regressions, the coefficient of the GB£/US$ *Lspread* is slightly greater than the EUR/US$ coefficient. In line with previous results, the coefficients of GB£/US$ and EUR/US$ *Lspread* are very close to each other and the coefficient of the CAD$/US$. *Lspread* is still smaller in the SURE estimation. Nevertheless, the gap between the most traded currency pairs and the CAD$/US$ is narrower when the SURE estimator is used. As a consequence, elasticities are the lowest when the contemporaneous correlations are taken into account.

Table 3 - Results of SURE estimations

	Constant		Lask		Lspread	
	Coeff.	p–value	Coeff.	p–value	Coeff.	p–value
EUR/US$	109.85	0.000	−419.39	0.000	−0.33	0.000
GBP/US$	155.55	0.000	−257.79	0.000	−0.36	0.000
CAD/US$	−96.50	0.000	541.62	0.000	−0.23	0.001
JPY/US$	947.39	0.000	−205.30	0.000	−0.08	0.145

5. PANEL DATA ANALYSIS

We next turn to the panel methodology. First, we test for the presence of a unit root in our series. Panel unit root tests proposed by Levin *et al*. (2002) and Im *et al*. (2003) are the most used tests in panel studies. However, the so-called *first generation* unit root tests (they assume cross sectional independence) are shown to be inconsistent in the presence of cross sectional dependence, because they suffer from severe size distortions (O'Connel, 1998; Philips and Sul, 2003; Banerjee *et al*., 2005). As exchange rates are correlated, the series in our panel (i.e. currency pairs) are likely to be pair-wise correlated. We thus test the cross section dependence in our panel data and we use the so-called *second generation* unit root tests.

5.1. Cross section dependence in panel data

Pesaran (2004) proposed a test for error cross section dependence which is an extension of the Breush and Pagan's test[10] (1980, see Peasaran, 2004). It is only valid for N relatively small and T sufficiently large. The residuals of the ADF regressions ($\hat{u}_{it}$) are used to compute the following Lagrange Multiplier (LM) test:

$$CD_{LM} = T\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}^{2}. \tag{7}$$

In equation (7), $\hat{\rho}_{ij}$ is the pair-wise correlation of residuals coefficient that is:

$$\hat{\rho}_{ij} = \hat{\rho}_{ji} = \frac{\sum_{t=1}^{T} e_{it}e_{jt}}{\left(\sum_{t=1}^{T} e_{it}^{2}\right)^{1/2}\left(\sum_{t=1}^{T} e_{jt}^{2}\right)^{1/2}},$$

where e_{it} is the ordinary least squares estimate of the residuals u_{it} in the ADF regression. Note that under the null hypothesis of no cross section dependence, the CD_{LM} statistic is asymptocally distributed as Chi-squared with $N(N-1)/2$ degrees of freedom.

Recognizing the shortcomings of the Breush Pagan's LM test when N is large, Pesaran proposed a simple alternative which is based on the pair-wise correlation coefficients and not on their squared one:

$$CD = \sqrt{\frac{2T}{N(N-1)}}\left(\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}\right), \tag{8}$$

where CD is $I.I.N(0,1)$.

When N is relatively small with respect to T, as in our panel, the Breush-Pagan test has good sample size properties. Nevertheless the Pesaran CD test is also performed for comparison. Table 4 reveals that the null hypothesis of no cross section dependence is rejected at 1% and 5% respectively. There is thus clearly strong dependence in our data. There are several potential origins for cross section dependence in the currency pairs' panel: one or more unobserved common factors, one or more observed common factors or a general form of cross sectional dependence. This result is undeniably not surprising, because, as mentioned earlier, all parities are expressed in relation to the dollar. In addition, there are many cross dependence factors which explain correlation between the exchange rates: economic policies disseminated by the Federal Reserve Bank or the European Central Bank (like the interest rates level) or economic indicators like the performance of the US economy (in November 2004 here). It is well known that some events or decisions from a country (notably the US) impact in a unexpected sense the other countries. Furthermore, the market psychology influences the trading and the traders react together to the same news or rumours.

10. "*Breush and Pagan proposed a Lagrange Multiplier (LM) statistic for testing the null hypothesis of zero equation error correlations*" (Pesaran, 2004).

Table 4 - Test for cross section dependence

	Empirical values	Critical values (5%)
CD_{LM} **statistic**	11.30	7.81
CD **statistic**	−3.21	−1.96

Note: trend and intercept and 60 lags are included.

5.2. Testing for unit root using second generation panel unit root tests

From the preceding tests outcomes, we note the presence of cross section dependence in our currency pairs. Thus, the second generation Panel Unit Root Tests (PURT hereafter) must be used in order to test for the order of our series. The literature on modelling of cross section dependence in panels is still developing (for recent surveys of this expanding literature, see Gutierrez, 2003; Hurlin and Mignon, 2005; Gegenbach, Palm and Urbain, 2006; Breitung and Pesaran, 2006; Baltagi and Pesaran 2007).

Three different panel tests allowing cross section dependence among the currency pairs are applied to our sample: Pesaran (2005, published in 2007), Moon and Perron (2004) and Bai and Ng (2004). These approaches make the assumption that the cross sectional dependence is induced by one or more common factors that vary along the time dimension, but are invariant across panel units.[11] TABLE 5 displays the results of those tests.

Considering the results of both CIPS and $P^c_{\hat{E}}$ test statistics, we conclude that all of our series are I(1), apart from *LVol*. Nevertheless, Moon and Perron tests indicate that all our variables are I(1) at 5% significance. Gegenbach, Palm and Urbain (2006) proposed a sketched approach of the different PURT: the pooled CIPS test has better power properties than the individual CAPS tests (not reproduced here), but the Moon and Perron tests (t^*_a and t^*_b) are more powerful than Pesaran CIPS tests. This analysis is in line with Gutierrez (2003) who considers that Moon and Perron tests show good size and power for different values of T and N. Consequently, we conclude that in an overall perspective, all of our series are I(1). Therefore, using in the next section the panel cointegration methods is warranted in our view.

Table 5 - Pooled panel unit root tests

Tests	**Pesaran**	**Moon and Perron (MP)**		**Bai and Ng (BN)**
Statistics	CIPS	t^*_a	t^*_b	$P^c_{\hat{E}}$
Lvol	−3.36 (0.01)	0.17 (0.57)	0.10 (0.54)	3.55 (0.00)
Lspread	−2.74 (0.09)	−36.43 (0.14)	32.86 (0.14)	0.67 (0.25)
Lask	−2.76 (0.08)	−1.42 (0.08)	−1.40 (0.08)	−0.96 (0.83)

Notes: 2 common factors, 60 lags and intercept are included in MP and BN tests.
P-values are listed between (.).

11. These recent tests have not yet been implemented in the usual econometric software like RATS, Eviews, SAS. However, it is possible to compute the second generation panel unit root tests using Matlab 7.1. The results of our panel unit root tests are presented in TABLE 4.

5.3. Testing for cointegration

We test for the existence of a cointegration relationship between *Lvol*, *Lask* and *Lspread* in our model (2). There are several panel cointegration tests: the single equations tests of Pedroni (1999, 2004) and Kao (1999) and the systems tests developed by Larsson, Lyhagen and Löthgren (2001). In this analysis, the test statistics of Pedroni are selected, because Gutierrez (2003) showed that these statistics have the best power among all the cointegration tests when the time dimension (T) is very high. Briefly, the single equations methods of Pedroni are panel extensions of the Engle and Granger (1987) method developed in section 2. In other words, the idea for the residual-based tests of Pedroni is to test for the existence of a unit root in the residuals of the spurious regression.

According to Pedroni (1999, 2004), the data generating processes (DGP) is:

$$y_{it} = \alpha_i + x'_{it}\beta_i + u_{it}, \quad N = 1, 2, 3, 4, \quad T = 1, \dots 1493, \tag{9}$$

where x'_{it} and β_i are $k \times 1$ vectors with $k = 2$ because there are two regressors *Lask* and *Lspread* in our model. Besides, y_{it} denotes *Lvol* and α_i is a fixed effect.

There are two different classes of statistics suggested by Pedroni (see table 1, 1999) among the seven statistics he proposed: four statistics are based on the within dimension (*panel* tests or *within*) and three on the between dimension (*group mean* tests or *between*). All these tests are based on the null hypothesis of no cointegration and unlike Kao (1999), some heterogeneity is introduced under the alternative hypothesis. Thus, there exists a cointegration relationship for each currency pair and this relationship is not necessarily the same for each currency pair.

The null hypothesis of no cointegration is given by $H0: \rho_i = 1$, where ρ is the first-order serial correlation coefficient of the residuals. The alternative hypothesis of stationary residuals (u_{it} in the equation (5)) in the *panel* tests is $H1: \rho_i = \rho < 1$, that is ρ is identical for $i = 1, \dots, N$. The *group* statistics are based on cross-section averages of individual estimators of ρ_i and they are specified against the alternative: $H1: \rho_i < 1$. Therefore, the *group* statistics are more general since they allow for heterogeneous coefficients under $H1$.

TABLE 6 points to the outcomes of the seven statistics proposed by Pedroni (1999, 2004). It is shown that the null hypothesis is rejected at 5% except for $Z_{vN,T}$. It is not a problem for many reasons: on the one hand, the null hypothesis of no cointegration is rejected for six of the considered statistics; on the other hand, Karaman Örsal (2008) argued with the help of Monte Carlo simulations that the *panel-t* test statistic ($Z^*_{tN,T}$) has the best size and size-adjusted power properties among all the Pedroni statistics. In accordance to Gutierrez (2003) the *group* ρ statistic ($\tilde{Z}_{\rho N,T-1}$) has the best power properties and in accordance to Wagner and Hlouskova (2007), the two tests of Pedroni applying the ADF principle perform test ($Z^*_{tN,T}$ and $\tilde{Z}^*_{tN,T}$) are the first choice.

Table 6 - Panel cointegration tests

Dimension	Statistic	Standardized values
Within	$Z_{vN,T}$	1.207 (0.114)
	$Z_{\rho N,T^{-1}}$	−3.903 (0.000)
	$Z_{tN,T}$	−10.336 (0.000)
	$Z^{*}_{tN,T}$	−5.873 (0.000)
Between	$\tilde{Z}_{\rho N,T^{-1}}$	−2.062 (0.020)
	$\tilde{Z}_{tN,T}$	−12.948 (0.000)
	$\tilde{Z}^{*}_{tN,T}$	−7.95 (0.000)

Note: Pedroni tests were computed using Rats code (PANCOINT).
P-values are in parentheses.

5.4. Estimating panel cointegration vector

Finally, we estimate the impact of the currency transaction tax *via* β_2:

$$Lvol_{it} = \beta_0 + \beta_1\ Lask_{it} + \beta_2\ Lspread_{it} + \alpha_i + u_{it}\,, \quad i = 1, \dots, N, \quad t = 1, \dots, T, \quad (10)$$

where α_i is a deterministic intercept for each currency pair *i*.

To obtain panel cointegrated estimates is not an easy task because the asymptotic properties of the estimators and the corresponding statistical tests are different from those of time series cointegration models. Indeed, equation (10) can be first estimated by ordinary least squares. (In this case, the estimator *Within* is used.) But in finite samples OLS estimator and *t*-statistic are biased. Therefore Phillips and Moon (1999) and Pedroni (2000) proposed a fully modified OLS estimator (FMOLS); Kao and Chiang (2000)[12] suggested a different approach based on the panel dynamic ordinary least squares (DOLS) estimator developed by Stock and Watson (1993) and Saikkonen (1991). These authors investigated, with Monte Carlo simulations, the finite sample properties of the OLS, FMOLS, and DOLS estimators. Their conclusions are the following ones:

- the OLS estimator presents a bias in finite samples; the OLS estimator, though superconsistent, would not be optimal for inference in this case;
- the FMOLS estimator does not improve over the OLS estimator in general;
- the DOLS estimator performs well in estimating cointegrated panel regressions.

Consequently we will use DOLS to estimate the model (10). What is peculiar to the DOLS estimator is that it includes lags and leads of the first difference to the set of cointegrating regressors in order to taken account the possible correlation of the error term. The DOLS estimates of the model (10) are reported in TABLE 7.

12. See also Mark and Sul (2003).

Table 7 - DOLS Fixed effects model estimation results

Variable	Coefficient	Standard error	t statistic	p-value
Constant	201.026	22.952	8.758	0.000
Lask	−147.648	16.176	−9.127	0.000
Lspread	−0.612	0.053	−11.428	0.000
N=4	NT=5972	$\bar{R}^2$=0.26	F=235.83	p-value (stat F)=0.000

Note: the choice of the correct lags and leads selection frequency is based on the Schwarz Bayesian information criterion as suggested by Westerlund (2005).

We derived several commentaries from the TABLE 7.

i) All regression coefficients are non null and thus significant.

ii) The constant value is positive, that is the foreign exchange trading volume is positive (201.026) whatever the *spread* and *ask* values. In addition, the fixed effects are −168.23, −186.15, −117.03 and 477.50 respectively. Once more, the yen currency pair leads to the more devious result. Considering the overall constant, it could be inferred that the effects of *Lspread* and *Lask* variables on the volume trading are less pronounced regarding the yen transactions.

iii) All the coefficients have the expected signs: an increase of the *ask* price or an increase in the *spread* has a negative impact on the foreign exchange volume for all currency pairs. Therefore, an increase of the transactions costs due to the tax introduction unambiguously leads to a decrease of the trading volume.

iv) The effects of *ask* price variations are stronger than *spread* variations. This finding is consistent with the foreign exchange market activity without Tobin tax, because the taxation is not the key decision variable for the traders.

v) Finally, the sharp value of the currency transaction tax elasticity is −0.61; in other words, an increase of 1% of the *spread* is corresponding to a decrease of more than half of the *forex* operations. This overall elasticity is higher (in absolute value) than the SURE elasticities outlined in the preceding section. This result is quite relevant because, considering the whole of the *forex* market, there are less opportunities of circumventing the tax and consequently the endogenous decreasing effect of the tax on the volume trading is more pronounced than in the preceding partial estimations.

We ensure to the robustness of our DOLS panel results by finally conducting some specification tests all based on the LR (Likelihood Ratio).

As outlined in the TABLE 8, the model (10) is quite robust. Indeed, the statistics values strongly reject the null hypothesis that the *Lask*, *Lspread* variables and fixed effects are redundant. Furthermore, the test statistic of omitted variable rejects the null that *Lask* and *Lspread* are jointly irrelevant; *Lbid* is not an omitted variable in the model (10).

Table 8 - Specification tests

Redundant variable				Omitted variable		Fixed effects	
Lask		Lspread		Lbid		$\alpha_1 = \alpha_2 = \alpha_3 = \alpha_4$	
LR Stat	p-value	LR Stat	p-value	LR Stat	p-value	LR Stat	p-value
76.25	0.000	208.87	0.000	8.124	0.004	1408.99	0.000

6. Concluding remarks

This paper is the first attempt to estimate the influence of a currency transaction tax on the foreign exchange market volume trading. The econometric estimations suggest that the *forex* trading volume could be significantly reduced by the Tobin tax. Nevertheless, the values obtained for the elasticites are heterogeneous with respect to the currency pairs and lead to some effects that are more pronounced than Frankel's previsions, but less marked than assessed in the institutional reports (French, Belgian, German reports). As we might expect, the most traded currency pairs are also the most traded exchange parities. Moreover, the elasticities values are the lowest when the SURE estimator is used, but the highest when the panel estimation is implemented.

We can first discuss the interpretation of those results in two ways. This significant decrease in the volume trading is good news for the Tobin tax advocates, who think that foreign exchange market transactions are largely of speculative nature. In this respect, our result would mean that the Tobin tax could curb the speculative trading volume as they expected.

In the opposite direction, it is a bad result for its opponents, since a high decreasing volume could have dangerous consequences for hedging and liquidity trades. In this way, a high rate Tobin Tax could destroy the existing market structure and the information efficiency of the market prices.

Finally, taking our results as a starting point, we can produce a new estimation (that is more accurate than the existing ones) of the potential revenues of such a Tobin tax. In this way, we assume that all major trading sites would agree to collect the tax. According to the last BIS survey data (BIS, 2007), global average daily net turnover, i.e. the global tax base was US$1880 billion. In accordance with Nissanke (2003) and the many official reports (French, Belgian, German…), it is also assumed that 10% and 20% of total turnover respectively are deducted from the tax base as non taxed trades and tax evasion. Furthermore, our estimates are based on the assumption that spreads are very low: 0.01-0.02% as suggested by Spahn (2002) and by our data set. Assuming that tax rates would be set at the very most to 0.02%, the currency transaction tax presents a potential of generating a maximum of US$40 billion for global multilateral projects.

To conclude this article, we suggest a set of extensions, which would improve the decline and the revenues estimates. Our results rely indeed on simplifying the assumption that all trades in the same currency pair have the same elasticity. The analysis would be further enriched if we could obtain more details about each realized transaction (liquidity trade or speculative trade, interbank liquidity trade or speculative hedge funds trade…). Because the foreign exchange market is very opaque, the achievement of this desirable extension will be a very difficult task.

F. B. & O. D.[13]

13. We are very grateful to Claude Diebolt, Bertrand Koebel, Phu Nguyen Van, Jacques Raynauld, Jimmy Lopez and two anonymous referees for helpful comments. An earlier version of this paper has benefited from comments of participants from GDR Monetary and Banking (2005), Journées de l'AFSE (2006), SCSE (2008) and SMYE (2008).

References

Aliber, R.Z., Chowdhry, B., Yan, S., 2003. Some evidence that a tax on foreign exchange transactions may increase volatility, *European Finance Review* 7 (3), 481-510.

Bai, J., Serena, Ng, 2004. A panic attack on unit roots and cointegration, *Econometrica* 72 (4), 1127-77.

Baltagi, B.H., (ed.), 2000. *Advances in Econometrics*, 15: *Nonstationary Panels, Panel Cointegration, and Dynamic Panels*, JAI Press, Amsterdam.

Baltagi, B.H., 2005. *Econometric Analysis of Panel Data*, Wiley, Chichester, 3rd edition.

Baltagi, B.H., Pesaran, M.H., 2007. Heterogeneity and cross section dependence in panel data models: Theory and applications introduction, *Journal of Applied Econometrics* 22 (2), 229-32.

Banerjee, A., Marcellino, M., Osbat, C., 2005. Testing for PPP: Should we use panel methods?, *Empirical Economics* 30 (1), May, 77-91.

Bird, G., Rajan, R.S., 2001. International currency taxation and currency stabilisation in developing countries, *The Journal of Development Studies* 37 (3), 21-38.

BIS (Bank for International Settlements), 2005. Triennal central bank survey. Foreign exchange and derivatives market activity 2004, Basel.

BIS (Bank for International Settlements), 2007. Triennal central bank survey. Foreign exchange and derivatives market activity 2007, Basel.

Bessembinder, H., 1994. Bid-ask spreads in the interbank foreign exchange markets, *Journal of Financial Economics* 35 (3), June, 317-48.

Bosco, B., Santoro, A., 2004. Tobin tax: A mean-variance approach, *Finanzarchiv* 60 (3), September, 446-59.

Breitung, J., 2000. The local power of some unit root tests for panel data, in Baltagi (ed), 161-78.

Breitung, J., Pesaran, M.H., 2005. Unit roots and cointegration in panels, *The Econometrics of Panel Data* (3rd Edition), Kluwer Academic Publishers, forthcoming.

Brown, R.L., Durbin, J., Evans, J.M., 1975. Techniques for testing the constancy of regression relationships over the time, *Journal of the Royal Statistical Society*, Series B, 37, 149-92.

Choi, I., 2001. Unit root tests for panel data, *Journal of International Money and Finance* 20 (2), April, 249-72.

Chaboud, A., Weinberg, S., 2002. Foreign exchange markets in the 1990s: Intraday market volatility and the growth of electronic trading, BIS Papers 12, August, 138-47.

Easley, D., O'Hara, M., 1992. Time and the process of security price adjustment, *Journal of Finance* 47 (2), June, 577-605.

Eichner, T., Wagener, A., 2005. The effects of the Tobin tax on the speculation and hedging, Discussion Paper, Hanover University, 1-13.

Elliott, G., Rothenberg, T.J., Stock, J.H., 1996. Efficient tests for an autoregressive unit root, *Econometrica* 64 (4), July, 813-36.

Engle, R.F., Granger, C.W.J., 1987. Co-integration and error correction: Representation, estimation and testing, *Econometrica* 55 (2), March, 251-76.

Felix, D., Sau, R., 1996. On the revenue potential and phasing in of the Tobin tax, in Grunberg *et al.* (eds), 223-54.

Frankel, J.A., 1996. How well do foreign exchange markets function might a Tobin tax help?, in Grunberg *et al.* (eds), 41-81.

Galati, G., 2000. Trading volumes, volatility and spreads in foreign exchange markets: Evidence from emerging market countries, BIS Working Papers 93, October.

Gaulier, G., Hurlin, C., Jean-Pierre, P., 1999. Testing convergence: A panel data approach, *Annales d'Economie et de Statistique* 55-56, 411-28.

Gengenbach, C., Palm, F., Urbain, J.P., 2006. Panel unit root tests in the presence of cross-sectional dependence: Comparisons and implications for modeling, Working Paper, University of Maastricht.

Glassman, D.A., 1987. Exchange rate risk and transaction costs: Evidence from bid-ask spreads, *Journal of International Money and Finance* 6 (4), December, 479-90.

Grunberg, I., Ul Haq, M., Kaul, I., (eds), 1996. *The Tobin Tax: Coping with Financial Volatility*, Oxford University Press, New York-Oxford.

Gutierrez, L., 2003. On the power of panel cointegration tests: a Monte Carlo comparison, *Economics Letters* 80 (1), July, 105-11.

Gutierrez, L., 2006. Panel unit-root tests for cross sectionally correlated panels: a Monte Carlo Comparison, *Oxford Bulletin of Economics and Statistics* 68 (4), August, 519-40.

Haberer, M., 2004. Might a securities transactions tax mitigate excess volatility? Some evidence from the literature, Discussion Paper 04/06, University of Konstanz.

Hartmann, P., 1998. *Currency Competition and Foreign Exchange Markets/the Dollar, the Yen and the Euro*, Cambridge University Press.

Hausman, J.A., 1978. Specification tests in econometrics, *Econometrica* 46 (6), November, 1251-71.

Hsiao, C., 1986. *Analysis of Panel Data*, Econometric Society Monographs 11, Cambridge University Press, Cambridge.

Hurlin, C., Mignon, V., 2005. Une synthèse des tests de racine unitaire sur données de panel, *Économie et Prévision* 169-170-171, 253-94.

Im, K.S., Pesaran, M.H., Shin, Y., 2003. Testing for unit roots in heterogeneous panels, *Journal of Econometrics* 115 (1), July, 53-74.

Johansen, S., 1988. Statistical analysis of cointegration vectors, *Journal of Economic Dynamics and Control* 12 (2/3), June-September, 231-54.

Johansen, S., 1995. *Likelihood-Based Inference in Cointegrated Vector Auto-Regressive Models*, Oxford University Press, Oxford.

Kao, C., 1999. Spurious regression and residual-based tests for cointegration in panel data, *Journal of Econometrics* 90 (5), 1-44.

Kao, C., Chiang, M.H., 2000. On the estimation and inference of a cointegrated regression in panel data, *Advances in Econometrics* 15, 179-222.

Levin, A., Lin, C.F., 1993. Unit root tests in panel data: New results, University of California at San Diego, Discussion Paper 93-56.

Levin, A., Lin, C.F., Chu, C.S.J., 2002. Unit root tests in panel data: Asymptotic and finite sample properties, *Journal of Econometrics* 108, 1-24.

McCoskey, S., Kao, C., 1998. A residual-based test of the null of cointegration in panel data, *Econometric Reviews* 17 (1), February, 57-84.

Moon, H.R., Perron, B., 2004. Testing for a unit root in panels with dynamic factors, *Journal of Econometrics* 122 (1), September, 81-126.

Newey, W.K., West, K.D., 1994. Automatic lag selection in covariance matrix estimation, *Review of Economic Studies* 61 (4), October, 631-53.

Nissanke, M., 2003. Revenue potential of the currency transaction tax for development finance: A critical appraisal, Wider Discussion Paper 2003/81, World Institute for Development Economic Research, the United Nations University, Helsinki.

O'Connell, P.G.J., 1998. The overvaluation of purchasing power parity, *Journal of International Economics* 44 (1), February, 1-19.

OCDE, 2002. Instabilité des marchés des changes et taxes sur les opérations financières, *Perspectives économiques de l'OCDE* 71, 221-34.

Pedroni, P., 1999. Critical values for cointegration tests in heterogeneous panels with multiple regressors, *Oxford Bulletin of Economics and Statistics* 61, 631-52.

Pedroni, P., 2000. Fully modified OLS for heterogeneous cointegrated panels, *Advances in Econometrics* 15, 93-130.

Pedroni, P., 2004. Panel cointegration: Asymptotic and finite sample properties of pooled time series tests with an application to the PPP hypothesis, *Econometric Theory* 20, 597-625.

Pesaran, M.H., 2003. A simple panel unit root test in the presence of cross section dependence, Department of Applied Economics, University of Cambridge, Cambridge Working Papers in Economics 0346.

Pesaran, M.H., 2004. General diagnostic tests for cross section dependence in panels, Department of Applied Economics, University of Cambridge, Cambridge Working Papers in Economics 0435.

Pesaran, M.H., 2007. A simple panel unit root test in the presence of cross section dependence, *Journal of Applied Econometrics* 22 (2), 265-312.

Phillips, P.C.B., 1987. Time series regression with a unit root, *Econometrica* 55 (2), March, 277-301.

Phillips, P.C.B., Moon, H.R., 1999. Linear regression limit theory for nonstationary panel data, *Econometrica* 67 (5), September, 1057-1111.

Phillips, P.C.B., Sul, D., 2003. Dynamic panel estimation and homogeneity testing under cross section dependence, *The Econometrics Journal* 6 (1), 217-59.

Phillips, P.C.B., Xiao, Z., 1998. A primer in unit root testing, *Journal of Economic Surveys* 12 (5), December, 423-69.

Saikkonen, P., 1991. Asymptotic efficient estimation of cointegrating regressions, *Econometric Theory 7* (1), March, 1-21.

Spahn, P.B., 2002. On the feasibility of a tax on foreign exchange transactions, Federal Ministry for Economic Cooperation and Development.

Stock, J.H., Watson, M.W., 1993. A simple estimator of cointegrating vectors in higher order integrated systems, *Econometrica* 61 (4), July, 783-820.

Tobin, J., 1974. *The New Economics One Decade Older*, Princeton University Press, Princeton, NJ.

Tobin, J., 1978. A proposal for monetary reform, *Eastern Economic Journal* IV (3-4), July-October, 153-59.

Westerlund, J., 2005. Data dependent endogeneity correction in cointegrated, *Oxford Bulletin of Economics and Statistics* 67(5), October, 691-705.

Yule, G.U., 1926. Why do we sometimes get non-sense correlations between time series? A study in sampling and the nature of time series, *Journal of the Royal Statistical Society* 89, 1-64.

Zellner, A., 1962. An efficient method of estimating seemingly unrelated regression equations and tests of aggregate bias, *Journal of the American Statistical Association* 57, 348-68.