Do Reuters Spreads Reflect Currencies' Differences in Global Trading Activity?

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DO REUTERS SPREADS REFLECT CURRENCIES' DIFFERENCES IN GLOBAL TRADING ACTIVITY ?

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Abstract: A new estimate of the long-run impact of trading activity on bid-ask spreads in the foreign exchange markets is realized with a short panel containing around-the-clock Reuters quotes and global transaction volumes. Individual and time effects are accounted for in an unbalanced random effects model. In accordance with liquidity effect explanations the volume parameter is found to have a negative sign, although not at a high level of significance. The volatility parameter is positive and strongly significant. While the structural parameters of the model appear to be stable over time, the residuals are groupwise heteroscedastic. The higher standard error in 1992 might reflect the dynamic developments in the world forex market since 1989. Reuters quoting (tick) frequency is also tested as a measure of trading activity in spread estimations. The results turn out to be very similar to those with trading volumes, in particular when an instrumental variable estimator is employed to account for measurement errors or possible endogeneity problems.

KEYWORDS: foreign exchange markets, bid-ask spreads, trading volumes, panel data, random effects models

JEL CLASSIFICATION: F31, G15, C33

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I. Introduction

It is economists' common sense that in relatively liquid financial markets transaction costs should be lower than in rather illiquid markets. However, some research has cast some doubt on this claim, in particular concerning the relation between trading activity and bid-ask spreads in the foreign exchange (forex) market. Glassman (1987) and Wei (1994) estimate a positive relation between daily forex trading volumes and spreads. With high frequency data Bollerslev and Domowitz (1993) and Lyons (1995) find spreads increasing with 'transaction frequency' or transaction size respectively.

This question of the effect of forex trading activity on spreads is interesting for at least three reasons. First, it is important for the discussion about market efficiency. Should the studies quoted above be right in general, then forex market growth might not be desirable from a welfare-theoretic point of view. Second, Krugman (1980), Black (1991), Hartmann (1994), and Rey (1997) argue that trading volume of a currency is an important factor for a currency to become a vehicle - a medium of exchange for currencies. Since their theoretical argument is based on a negative impact of volume on transaction costs, as measured by bid-ask spreads, the current theory on vehicle currencies in the forex market would have to be rewritten, if the (long-run) relationship between forex volumes and spreads turns out to be positive. Finally, the relationship between trading volumes and bid or ask prices may be relevant for (short-run) exchange-rate forecasting.

The theoretical finance literature derived some arguments when dealer spreads could go up with transaction volume. Copeland and Galai (1983) and Glosten and Milgrom (1985) develop the information cost model of the bid-ask spread in financial markets. In this framework, if dealers perceive transaction volume to be positively correlated with the probability of getting into a transaction with a better informed counterpart, then higher volume increases their expected costs of making a market, which has to be offset by a higher spread to deter some of the informed traders or increase earnings from uninformed (liquidity-motivated) traders.[1] Wei (1994) points out that such a positive impact might prevail in the short run, while in the long run the correlation should be negative. This claim can be based on economies of scale in market making, as

[1] For a rich discussion of the subtleties of the informational content of forex transaction volume, see Lyons (1995, 1996). In some inventory cost and information cost models bid-ask spreads widen with the *size* of transactions (Stoll, 1978; Lyons, 1995). Glassman (1987) argues that a positive impact of volumes on spreads could result if trader disagreement drives transactions.

highlighted in order processing cost models of bid-ask spreads, like Black (1991), but also the inventory cost model of Stoll (1978). Another justification for a negative long-run correlation would be thick-market externalities in a search cost framework (Chrystal, 1984). It is further substantiated by recent empirical findings showing that the permanent (predictable) component of *daily* volume changes decreases spreads, while the transitory (unpredictable) component - reflecting information arrival - increases them (Bessembinder, 1994; Hartmann, 1996; Jorion, 1996).

In the present paper, in the first place, a new attempt to estimate the *long-run* volume-spread relation is undertaken using a short panel data set of monthly forex turnovers in many different currency pairs. In line with theory the long-run volume effect is consistently estimated to be negative. However, even after meticulous adjustment for the panel characteristics of the data it is only weakly significant (usually between 5 and 10 percent for a two-sided test). Second, the quality of Reuters FXFX tick frequency as a proxy for trading volume is evaluated. In fact, we can show in an auxiliary regression that monthly Reuters ticks are quite strongly correlated with monthly trading volumes in our sample, although the relationship found is unstable over time. Instrumental variable spread estimations with either measure of trading activity (volumes or ticks) lead to similar results. Hence, while FXFX ticks seem to be an imprecise measure of trading activity for high-frequency estimations (Goodhart et al., 1996), they turn out to perform quite well for more long-run analyses.

The next section discusses data problems related to volume-spread estimations in the forex market and reviews the empirical literature. Because of the general scarcity of information about trading volumes in this market, particular emphasis is given on the availability and quality of different types of volume data. After a brief description of the econometric approach in the following section, section IV summarizes the results found on the spread-volume relationship and section V those found on the volume-ticks and spread-ticks relationships. Conclusions and implications for future research are contained in the final section. A more comprehensive description of the econometric techniques applied are put in an appendix.

II. Data and Measurement Issues

Recent theories of forex bid-ask spreads start out from the rationale of a single dealer (market-maker) who determines his buying and selling rates (bid and ask rates) on the basis on his evaluation of the current situation in the whole market. For example, Black's (1991) order processing cost model leads to a dealer spread of the following simplified form

Install Equation Editor and double-click here to view equation. **1**

where s_{ij} denotes a dealer's fractional spread (ask rate minus bid rate divided by the middle rate) for the currency pair ij, σ_{ij} his expectation on the exchange rate volatility (at given order flow), and v_{ij} the expected order flow of currency i against currency j. (α is a parameter.)[2] These models predict, inter alia, that for a single forex dealer the percentage spreads of two currency pairs differ inversely to the dealer's respective order flow in these two currency pairs (at constant volatility). Allowing for volume and volatility elasticities of the spread (β_1, β_2) different from unity this equation might be rewritten in logs.

Install Equation Editor and double-click here to view equation. **2**

(2) and many other forex micro-structure models of the bid-ask spread are not specified in a way which corresponds directly to the available data. In particular, because of the fragmentalized nature of the forex market and the dealers' interest that their competitors do not know their positions, satisfying volume data in general are difficult and on the level of single dealers almost impossible to obtain. In addition, those volume data which are available for researchers usually come at a much lower frequencies (daily or monthly) than the actual updating of spreads in the market. However, using some simple algebra and the assumption that β_1 does not differ between dealers active in ij one can easily show from (2) that the *average market spread*over several periods and dealers is a function of *aggregate market turnover*

1. Measurement of trading volumes

Econometricians estimating the effect of trading volumes on bid-ask spreads in the forex market have used five different sources to measure trading activity. These are forex turnovers as collected by central banks in the main trading centers, forex futures turnovers at the Chicago International Monetary Market (IMM), Tokyo forex broker turnovers, quoting frequency on the Reuters FXFX page, and single banks' transaction

[2] A similar formula for the spread was derived already by Müller (1986).

volumes.

a) BIS global volumes

In the early 1980s the New York Fed, the Bank of England and the Bank of Japan began collecting forex turnovers as reported by most dealers (and brokers) in their respective markets for a particular month (usually April). Since 1989 the BIS coordinates this survey for a much larger number of central banks (21 in 1989 - excluding Germany's Bundesbank -, and 26 in 1992). The BIS surveys offer different breakdowns of volumes, for example according to trading centers, currency pairs, counterparties (inter-dealer versus dealer-customer) and contract types (spot, forward etc.). Because of almost complete coverage in most markets local double-counting, arising from the fact that each transaction is reported by two forex dealers, can be corrected for relatively precisely. However, since the national reports do not break down cross-border volume according to counterpart countries imprecise corrections for cross-border double-counting enter some errors. A second problem with the BIS turnover surveys is their extremely low frequency. Since they are relatively costly for both private banks and central banks, they are undertaken only every three years, which precludes any time series techniques. Finally, one might object that monthly market volumes do not correspond precisely to the short-term expected individual-dealer turnovers as they enter spread equations like (1).

The branch of the volume-spread literature exploiting these three-annually surveys was pioneered by Black (1991), who undertook a pooling regression of 4 annual observations (1980, 1983, 1986, 1989) for 7 dollar markets. He finds a significantly positive sign for a composite variable, where exchange rate volatility enters in the numerator and trading volume in the denominator. A simple cross-section regression by Bingham (1991) with 20 observations from the 1989 BIS report resulted in a negative, but insignificant volume parameter. Because of their large coverage and relatively deep breakdowns we take the BIS volumes up again for the present study. However, we try to improve several shortcomings of the previous papers, in particular concerning the econometric method, the number of observations, and the measurement of spreads as well as volumes.

More precisely, we use global spot inter-dealer turnovers for 22 currency pairs over April 1989 and 33 currency pairs over April 1992. This volume measure is more exact than those employed in the former studies with BIS data which take *total* volumes, containing also swap and forward as well as dealer-customer transactions. Each observation is adjusted for differences in the number of business days in financial centres and therefore reflects volume for a 'representative' business day for that month.[3] Part of the observations could be taken directly from the BIS surveys (BIS 1989, 1992), others are computed from the national surveys sent to us by central banks. In the latter case global turnover of a currency pair was approximated by the sum of the *local* turnovers in each currency's domestic trading center plus the higher one of the two *cross-border* turnover numbers reported for each center. (This convention may induce a slight downward bias in the measurement of those currrencies' volumes.) In a few cases the share of spot inter-dealer transactions in overall volume had to be approximated. Table 1 reports the data for April 1992 in descending order.

[3]Hence, even though daily representative volumes they rather reflect *monthly* trading activity (see BIS, 1993). The total number of 55 observations was limited by the number of rates quoted in Reuters with which the volume data have to be matched.

Table 1: Reuters Spreads, Trading Volumes, Quoting Frequencies, and Exchange RateVolatilities in April 1992

Market (1)	Spread (2)	Volume (3)	Ticks (4)	Volatility (5)
DEM/USD	45.615	87938.5	112184	313.52
JPY/USD	57.107	44147.8	51565	223.24
USD/GBP	46.236	22876	39979	276.93
CHF/USD	61.811	16736.9	44231	364.47
DEM/GBP	41.383	15712.3	6617	182.63
JPY/DEM	41.407	12298.7	7123	305.05
CHF/DEM	34.045	9110.2	6472	160.53
FRF/DEM	11.114	6701.1	5476	40.47
CAD/USD	42.599	5107.3	10072	142.18
DEM/XEU	24.347	5072.1	4629	40.9
USD/AUD	67.079	3682.3	20879	205.51
ITL/USD	59.812	3398.5	8892	282.53
ITL/DEM	57.053	3293.5	3984	62.16
FRF/USD	37.715	3023.4	28414	315.2
HKD/USD	13.784	2623.2	1085	28.06
NLG/DEM	8.462	1944.9	5101	24.29
USD/XEU	54.857	1383.4	10666	306.08
SGD/USD	63.679	1191.6	3053	134.55
NLG/USD	36.427	1135.8	18874	318.1
ESP/USD	51.384	850.5	5554	314.29
FIM/USD	46.698	745	9053	468.9
SEK/USD	33.573	680.7	54866	339.17
ZAR/USD	61.86	473.4	710	155.18
SAR/USD	8.79	377.1	206	2.06
DKK/USD	45.281	333.5	2048	304.68
USD/NZD	131.307	270.9	6046	236.85
GRD/USD	44.025	251.3	1103	269.57
ATS/USD	43.998	220.3	796	326.48
BEF/USD	35.684	186	2885	340.08
MYR/USD	41.189	154.5	4103	154.71
NOK/USD	34.126	118.1	7518	274.33
USD/IEP	61.236	48.3	743	326.28
PTE/USD	124.542	45	20381	394.06

b) IMM futures volumes

Daily currency futures volumes at the Chicago International Monetary Market are readily available since the number of traded contracts per day are reported by the market and the sizes of contracts are standardized. The obvious disadvantage of these data is that they represent only a small share of total (about 1 percent) as well as forward forex trading. (While Bessembinder (1994) argues that the stock-market experience shows that the correlation between spot and futures volumes is relatively high, Hartmann (1996) points to differences between stochastic processes fitted on forex futures and forex spot volumes.) Furthermore, while available time-series are quite long the IMM provides markets only for the six *major* currencies against the US dollar.

In her seminal paper Glassman (1987) matched these daily volumes with the corresponding futures price data. She estimated six spread equations with seemingly unrelated regression (SUR) for 1975 through 1983. In three cases a *positive* impact of volume on spreads is significant at the 5-percent (or lower) level, in two cases it is also positive but on a higher level of significance, and in the remaining case a negative coefficient is completely insignificant.

Recently Bessembinder (1994) tested a model by Easley and O'Hara (1992) with the futures data, suggesting that expected and unexpected volume should have opposite effects on spreads. The former volume component should reduce spreads through the order processing cost channel, while the latter should increase spreads through the information cost channel. Bessembinder does feasible generalized least squares (FGLS) estimations for 4 dollar markets between January 1979 and December 1992. He fits a stochastic process on the volume series in order to distinguish between the 'expected' and the 'unexpected' component. The parameters estimated for 'expected' volume are consistently negative, while those for 'unexpected' volume are always positive. However, the significance levels of these parameter estimates are sensitive to the currency regarded and to the convention on the measurement of spreads. In fact, the only cases where the hypothesis of the volume effect being zero can be rejected on a significance level of 5 percent (or better) are for spreads quoted in European terms in the markets for dollar/pound and dollar/yen ('expected' volume), as well as dollar/mark ('unexpected' volume).

Jorion (1996), applying a similar decomposition as Bessembinder for seven years of daily dollar/mark futures volumes coupled with spot price data (1985 through 1992), estimates (heteroscedasticity-consistent OLS) a negative coefficient for 'expected' volume (significance level 5 percent) and an insignificant (positive) coefficient for 'unexpected' volume.

c) Tokyo broker volumes

Only very recently a longer time series of *spot* forex volumes has been discovered (Wei, 1994; Hartmann, 1996). These data are published by the financial newspaper *Nihon Keizai Shimbun*, for the dollar/yen market exclusively. In Tokyo all forex *brokers* have to report their volume of transactions in dollar/yen concluded between opening and 3.30pm (local time) to the Bank of Japan. Although dollar/yen is traded world-wide, the Japanese part is a good proxy for global spot forex turnover in this currency pair.

Nevertheless, this time series has also some drawbacks. It might be affected by changes in the share of brokered deals in total trading, as has been the case in Japan (Bank of Japan, 1993). Also, broker volume might still be slightly different from direct inter-dealer volume, for example containing larger single transaction sizes.[4] Nonetheless, it certainly is a better proxy of global spot turnover than futures volume and it comes at a higher frequency than BIS survey data.

Wei (1994), whose main interest is in volatilities and spreads though, uses the Tokyo broker volumes in an univariate ordinary least squares (OLS) regression with monthly data (only one trading day per month) between 1983 and 1990 in order to estimate the volume effect for dollar/yen alone. Just as Glassman (1987) he finds the volume parameter to be positive, both in levels and in logs, although it is never more significant than 10 percent. In contrast, Hartmann (1996) fully exploits the daily frequency of these data for the period of 1987 through 1994 and further elaborates on Bessembinder's (1994) methodology. In particular, he accounts for the endogeneity of unpredictable turnover by introducing unpredictable Reuters FXFX tick frequency (see d) below) as an instrumental variable. With these improvements Bessembinder's qualitative results are confirmed, but the results turn out to be statistically much stronger, both for the negative 'expected' volume parameter (5 percent significance) and the positive 'unexpected' volume parameter (1 percent significance).

[4] See Hartmann (1996) for a full discussion of the Tokyo broker volumes.

Do these findings imply that unpredictable volume is an omitted variable in our estimations below? The answer is no. Bessembinder (1994), Hartmann (1996) and Jorion (1996) model daily unpredictable volumes as the residuals of AR(I)MA processes, which by definition have zero mean. Apart from some unsystematic error they cancel out when aggregated over longer time periods such as a month. BIS trading volumes therefore measure *predictable* (long-run) turnover alone, which in conjunction with the results of the three previous studies implies a strong prior in favor of a negative volume parameter in our spread estimations.

d) Reuters tick frequency

The use of quoting (tick) frequency as a proxy for trading volume (or market activity) was pioneered by Demos and Goodhart (1992) and Bollerslev and Domowitz (1993). In the interbank forex market practically all participants are connected with the Reuters information system where dealing banks feed in their bid and ask rates. The quotes are continuous 24 hours a day, 7 days a week. The obvious advantage of Reuters ticks is their extremely high frequency, which allows for measures of activity with time horizons much closer to those likely to predominate in the real market. Additionally, the quoting bank and its location can be identified from the respective Reuters page.

On the other hand, there are a number of disadvantages. First, the quoting frequency of dealers might not always be a precise measure of their trading activity.[5] It is not clear whether real transactions are done at the quoted prices and, if yes, at which amounts. While in normal times it can be expected that quoting frequency exceeds transactions frequency, in hectic market situations it may be the other way round, because dealers are too busy to feed in new quotes.

Second, the huge amount of data which accumulates rapidly when the Reuters information is stored requires powerful computer facilities, including automatic filters to clean the data from outliers, for example wrongly feeded quotes. This might be the main reason why Goodhart's data set, used in the two studies quoted above, is limited to less than three months. The forex consultancy firm Olsen & Associates (Zurich) has such

[5] As an extreme example, it happens sometimes that a major dealer in a small, relatively illiquid market quotes in short time intervals, quasi automatically 24 hours a day. Although it is Reuters' policy to prevent such behavior, sometimes a bank trying to advertise its presence in this market succeeds in doing this for weeks. The last entry (PTE/USD) in the third culomn of table 1 illustrates the point. In this particular case almost all quotes came from the same bank.

facilities as its disposal and stores Reuters quotes since the middle of the 1980s (see Dacorogna et al., 1993).

Demos and Goodhart (1992) use plain tick frequency (on Reuters page FXFX) as a measure of half-hourly market activity in dollar/mark and dollar/yen between April and June 1989. However, with a trivariate VAR estimation (also including volatilities) they do not find any significant correlations between ticks and spreads. Bollerslev and Domowitz (1993) also exploit Goodhart's data set, but - in order to correct for some of the problems mentioned above - simulate synthetic dollar/mark 'transaction frequencies' counting one transaction whenever two banks' spreads overlap within a five-minute time interval. A maximum-likelihood estimation of a GARCH model shows the impact of the length of time between two 'trades' on the spread to be significantly negative (the number of plain quote arrivals as another measure of market activity is clearly insignificant).[6] Davé (1993) - who has full access to the O&A data base - plots hourly dollar/mark Reuters spreads for January 1986 through September 1993 against hourly quoting frequency and hourly price changes. In contrast to the former studies his graphic shows a clear trade-off between spread size and 'market activity' *at constant volatility*.

Hartmann (1996) argues that tick frequency might be a measure of the rate of information arrival over the trading day, which is the mixing variable in models of the mixture-of-distributions hypothesis such as Tauchen and Pitts (1983) driving unpredictable trading volume. After decomposing daily ticks in a predictable and an unpredictable component, he shows that the latter performs very well as an instrumental variable for *unpredictable* dollar/yen spot turnover (see c) above). In another high-frequency analysis Lyons (1996) finds that the informational content of transaction size is strong in times of high quoting frequency (in the Reuters 2000-1 system), while it is weak in times of low quoting frequency (see also e) below).

However, Goodhart et al. (1996), by comparing one day of Reuters FXFX tick frequencies with *transactions data* from the trading system Reuters D2000-2, have thrown some doubt on the quality of the former as a proxy for trading activity in high-frequency analysis. In section V we test whether differences in *monthly* tick frequencies between bilateral markets reflect differences in global trading volumes of these markets and whether they can be successfully applied as a measure of *predictable* trading activity in more long-run spread

[6] It should be emphasized though that the GARCH model has not turned out to be a good econometric specification for *intra-day* foreign exchange data (Andersen and Bollerslev, 1994; Guillaume et al., 1995). Also, given the imprecision in the timing of FXFX quotes, a five-minute sampling interval looks extremely short.

estimations.

e) Individual dealer transactions volumes

Since most banks are extremely concerned about revealing their forex positions to their competitors access to real transaction order flows of single dealers has been extremely rare.[7] Nevertheless, Lyons (1995, 1996) has five days of transactions data (Reuters 2000-1 system) for one US dealer and one US broker for dollar/mark in August 1992. The originality of Lyons' analysis nonwithstanding his data set has the problem of rather limited coverage, both with respect to time and with respect to the number of dealers covered. (The week considered is just in the run up of the 1992 EMS crisis.)

Although Lyons (1995) focuses more on the explanation of intra-day volatility his econometric test also implies that - in line with information cost theories - dealer spreads widen with the *size* of single transactions. However, in a follow-up project (Lyons, 1996) evidence is provided that transaction size can be *less* informative when *transaction* frequency is high than when it is low. This contrasts with his result on quoting frequency as a measure of trading activity (see d) above).

Summing up this survey, one might conclude that completely satisfying forex volume data for the test of the hypotheses we are interested in do not exist at the present time, and that any analysis of it will therefore remain imperfect in some way. Nonetheless, it seems interesting to check whether the time-series results of Bessembinder (1994), Hartmann (1996), and Jorion (1996) on the long-run effect of trading volumes on bid-ask spreads can be confirmed with monthly cross-sectional data.

2. Measurement of bid-ask spreads

Surprisingly, bid and ask prices in the foreign exchange market also pose some measurement problems. The main reason is that most available data are *quoted* prices rather than real transaction prices. This is reflected in the spread-volume literature. Glassman (1987) uses daily quotes of a single Chicago futures dealer. Demos and Goodhart (1992), Bollerslev and Domowitz (1993), as well as Davé (1993) have access to

[7] Goodhart and Guigale (1988) have daily trading volumes from two London dealers. Unfortunately this data set, which covered several months in 1986, seems to be lost.

continuous Reuters quotes (page FXFX). Bingham (1991), Bessembinder (1994), and Jorion (1996) use daily data from the DRI data bank, which are Reuters quotes of a 'representative' dealing bank at some time during the day (e.g. at London closing). Black (1991) exploits fixing rates in some European markets as published by the Bundesbank.

In this study we use monthly tick-wise averages of quoted relative Reuters spreads (pages FXFX and WXWY).[8]

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(L is the number of Reuters ticks during a month for currency pair ij, a and b indicate ask and bid prices.)

These Reuters data have some peculiar features. First, quoted spreads are usually larger than traded spreads. (For example, it can be estimated from table 1 and Lyons (1995) that USD/DEM quoted spreads are about two to three times larger.) Second, the distribution of absolute (difference between ask and bid rate) Reuters spreads is discrete, with most of the mass on only a few numbers, such as 5, 7, and 10 basis points for dollar/mark (Goodhart and Curcio (1991), Bollerslev and Melvin (1994)), although with relative (or fractional) spreads this pattern becomes blurred. Both features indicate that the quoted spreads are less variable than the traded spreads and therefore suffer from lost information. However, Bollerslev and Melvin (1994) show that there still is a relatively high degree of variability in continuously quoted absolute Reuters spreads. Of course, daily 'representative' quotes are even less informative. Absolute fixing quotes hardly move over time. This might justify the use of all Reuters spreads quoted over a month as a proxy for transaction costs in the forex market.

3. Measurement of volatilities

Volatility is measured as the monthly average of daily absolute changes of Reuters middle rates.

4

[8] All exchange rate data in this paper were kindly provided by Olsen & Associates (Zurich).

where

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is the middle rate and D is the total number of days for a month.

Daily averages are chosen to make the measure compatible with the BIS volume figures described above. Absolute price changes are given preference to squared price changes, because of the absence of fourth moments in the distribution of exchange rate returns (Guillaume et al., forthcoming). This choice was taken for reasons of precision, but it does not affect our qualitative results.[9] As in the case of volumes these monthly averages rather reflect the aggregate of *predictable* price changes over shorter time intervals.

Wei (1994) and Jorion (1996) extract daily expected exchange rate volatilities from forex option prices as quoted at Chicago's IMM or the Philadelphia Exchange. Notice that implied volatilities cannot be used in our cross-section approach, since option prices are observable only for a small subset of bilateral markets over the period considered. Currency options for medium-range and small markets (see table 1), like e.g. SGD/USD, ATS/USD, DEM/XEU (XEU=ecu), and many others, are traded over-the-counter and also relatively illiquid.

III. Econometric Strategy

Our data set consists of an unbalanced panel with two periods, 22 observations in period 1 (April 1989) and 33 observations in period 2 (April 1992). In order to estimate the impact of trading volumes on bid-ask spreads we apply a variety of panel techniques with relative spreads (y) as the dependent variable and trading volumes (x_1) as well as exchange rate volatilities (x_2) as the independent variables (all variables in logs as specified before). Simple regressions are run for both periods separately and for the pooled sample (total

[9] Analyses of the tails of exchange rate return distributions have discovered the non-existence of fourth moments. For example, it has been shown by examinations of the autocorrelation functions of different volatility measures that the structure of volatility is better captured by absolute returns than by squared returns (Guillaume et al., forthcoming). Notice that the computation of the OLS and FGLS estimators with squared returns as explanatory variables would involve the inversion of returns raised to the power of 4.

model). Then spreads are estimated again with one-way and two-way error components specifications, including within and between transformations as well as random effects models.

The two-way random effects model as the most general specification can be written as

Install Equation Editor and double-click here to view equation. **6**

μ_i denotes an unobservable individual effect for currency pair i and λ_t an unobservable time effect for period t. Both are defined as stochastic deviations from the general intercept β_0, having zero means and standard deviations σ_μ and σ_λ. ε_{it} is the remaining error. All error components are assumed to be orthogonal to each other and to the observable explanatory variables. Individual and time effects are introduced in addition to the observable explanatory variables to avoid the potentially distortionary effects of omitted variables, such as differences in market micro-structure or market-wide technical progress.

The one-way random effects model is analogous, except that time effects are assumed to be negligible (λ_t=0). In both cases the covariance matrix of the errors is non-diagonal requiring the application of a feasible generalized least squares estimator (FGLS). For unbalanced panels the FGLS estimation of the two-way model is much harder to programme than that of the one-way specification, which might partly explain the latter's popularity among applied panel econometricians. Also, given that most panel data sets are still relatively short, little may be gained by the introduction of time effects. To be on the safe side, we estimate both specifications - the one-way model with the FGLS estimator described in Greene (1993) and the two-way model with the FGLS esimator proposed by Wansbeek and Kapteyn (1989).

In both specificiations we test for the presence of stochastic individual/time effects and for the absence of 'Mundlak bias'. The latter refers to the potential inconsistency of the FGLS estimators if the omitted variables, captured by the random effects, are correlated with the observable explanatory variables resulting in non-orthogonality of the regression residuals η and the system matrix X (Mundlak, 1978). For the one-way model we adjust the F-tests for individual effects (H_0: σ_μ=0) and for 'Mundlak bias' (H_0: $E(\mu|X)\neq0$) depicted by Dormont (1989) to the case of unbalanced panels. For the two-way model we apply the Honda (1985) and Gourieroux-Holly-Montfort (1982) tests for random effects (H_0: $\sigma_\mu=\sigma_\lambda=0$) and the Hausman (1978) test for Mundlak's omitted variable bias (H_0: $E(\mu|X)\neq0$ and $E(\lambda|X)\neq0$).

In the absence of inconsistency these random effects specifications are preferable to within transformations (or fixed effects specifications), where individual and time effects are filtered out through some weighting procedure (or captured by dummy variables), because of the important efficiency losses when applying the latter (Taylor, 1980; Baltagi, 1995). However, for completeness we also report within estimates below. More detailed descriptions of all these techniques can be found in the appendix.

IV. Spreads and Trading Volumes

Our results when applying the panel estimation techniques described in the preceding section are summarized in table 2. All results reported there are for log-linear equations such as (2). In fact, Box-Cox transformations, not reported here, suggested that a log-specification fits the data better, and additionally

1989 (OLS)	1.817 (5.93, 9.34)	-0.034 (-1.21, -1.48)	0.415 (8.02, 15,24)	0.75	0.23
1992 (OLS)	1.810 (3.92, 4.86)	-0.031 (-0.81, -0.98)	0.414 (6.21, 6.67)	0.54	0.42
Total (OLS)	1.817 (7.35, 9.89)	-0.033 (-1.52, -1.72)	0.414 (10.53, 13.72)	0.94	0.97
One-way model					
Between (OLS)	1.789 (4.89, 6.77)	-0.039 (-1.23, -1.68)	0.430 (7.62, 9.23)	0.85	1.01
Within (OLS)	1.983 (4.47, 5.37)	-0.048 (-1.26, -0.96)	0.406 (4.88, 7.06)	0.99	0.33
Random effects (FGLS)	1.834 (6.97, 8.50)	-0.038 (-1.68, -1.72)	0.420 (9.92, 12.26)	0.97	0.65
Tests [p-value]	Indiv. eff.: $F_{19,19}$=3.156 [0.01]		Mundlak test: $F_{3,49}$=0.018 [1.00]		Hausman test
Two-way model					
Within (OLS)	1.983 (4.64, 7.89)	-0.048 (-0.83, -1.41)	0.406 (6.10, 10.37)	0.99	0.29
Random effects (FGLS)	1.866 (6.98, 7.05)	-0.042 (-1.74, -1.88)	0.419 (9.07,10.62)	0.95	0.95
Tests [p-value]	Honda test: N(0,1)=2.342 [0.02]		GHM test: χ^2(mix)=3.586 [0.08 (by lin. extrapol.)		

Table 3: Analyses of covariance for spread estimations

Model	Chow test [p-value]	Test on const. intercept, other parameters const. [p-value]	Test on const. volume parameter, other para const. [p-value]
One-way model with volumes (FGLS)	$F_{3,49}$=0.017 [1.00]	$F_{1,51}$=0.044 [0.83]	$F_{1,51}$=0.048 [0.83]
Two-way model with volumes (within, OLS)	$F_{3,15}$=0.000 [1.00]	---	$F_{1,18}$=0.102 [0.75]
One-way model with ticks (FGLS)	$F_{3,46}$=0.083 [0.97]	$F_{1,48}$=0.005 [0.94]	$F_{1,48}$=0.001 [0.98]
Two-way model with ticks (within, OLS)	$F_{3,12}$=0.000 [1.00]	---	$F_{1,15}$=0.119 [0.73]

Third, there is evidence of both stochastic individual and time effects (table 2). Assuming no time effects (one-way model) the F-test on the absence of individual effects strongly rejects the null hypothesis (level of significance below 1 percent). For the two-way model the joint hypothesis of the absence of individual as well as time effects is rejected at 2 and 8 percent respectively, depending on the test chosen. (We have also tested for the absence of time effects, assuming no individual effects. In this case, which is not reported here, the null could be weakly rejected.) Moreover, 'Mundlak' and Hausman tests cannot reject the hypothesis of no omitted variable bias. This finding indicates that the ignorance of individual and time effects in former studies might not have caused biased parameter estimates but 'only' efficiency losses and biased inference.

What is the economic intuition behind these time and individual effects? As regards to the former, technical progress for example may reduce spreads for *all* markets from one period to the other. As can be seen from the separate regressions at the top of table 2, the intercept in 1992 is slightly lower than that in 1989. However, the difference is very small and the structural stability analysis in table 3 shows that it is, in fact, statistically not significant. One interpretation is that the *realization* of the random time effect has been quite similar in 1989 and 1992. Individual effects are stronger in our sample. (A fixed effects specification with dummies for bilateral markets would visualize them, but is not reported to save space.) This cross-sectional variation of bid-ask spreads may reflect - apart from measurement errors in (expected) volumes and volatilities - differences in the intensity of competition and in the micro-structure of bilateral markets (unrelated to volume or volatility). Another explanation might be differences in the coverage of market participants feeding their quotes in the Reuters system. Finally, exchange rate regimes or capital controls might play a role.

In order to avoid any biases in the *inference* about parameter significance we should chose the two-way random effects model as the most efficient consistent specification. When interpreting the parameter estimates found, we will therefore concentrate on the results reported at the bottom of the table. However, as could be expected from the short time dimension of our panel and the relative weakness of the time effects as compared to the individual effects, parameter estimates as well as standard errors differ only marginally from those found for the one-way random effects model further up.

The next result, already found in numerous studies before, is that the impact of exchange rate volatility on forex spreads is positive and strongly significant (significance level below 1 percent).[10] As expected the volume parameter is negative, although at a much lower level of significance. The two-tailed t-test of the hypothesis that it is equal to zero rejects the null at a significance level just below 9 or 7 percent with usual or heteroscedasticity-consistent (White) standard errors. (In any case, neither a White nor a Breusch-Pagan heteroscedasticity test comes close to rejecting the null hypothesis of homoscedastic residuals.) A one-tailed t-test of the null that the volume parameter is greater than or equal to zero rejects at below 5 percent in both cases. The conclusion is that the effect of monthly trading volume on Reuters spreads in our sample seems to be negative. However, the level of statistical significance is still uncomfortably low. If one compares the sample size available here (55) to that now available for daily time series estimations (almost 2000), then it is not surprising that the effect does not come out as strong as in the daily regressions (Hartmann, 1996). Important is that the long-run approach chosen here and the short-run approach with volume decomposition lead to perfectly compatible results.[11]

In order to elaborate further on the source of the individual effects which came out so strongly from the previous regressions, we introduced two other explanatory variables, a dummy for fixed exchange rate regimes prevailing in bilateral markets and another dummy for the presence of capital controls for at least one of the currencies in a bilateral market.[12] With these extensions the broad picture remained the same as before, in particular concerning the adequacy of the random effects specification. The absence of individual effects was still rejected, although - indeed - at a reduced level of significance. As could be expected, the exchange regime parameter had a negative and the capital control parameter a positive sign, but since both were clearly insignificant in the final specification and also produced multicollinearity problems we abstain from reporting further details.[13]

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[10]This result was already found by Agmon and Barnea (1977), Bingham (1991), Black (1991), Bollerslev and Melvin (1994), Boothe (1988), Glassman (1987).

[11]We also examined the residuals of these estimations and did not find indications for outliers in the data which might have influenced unduly the results found.

[12]The relevant information was extracted from the International Monetary Fund's report on exchange arrangements and exchange restrictions (IMF, 1990, 1993).

[13]For example fixed exchange rate regimes will usually result in lower exchange rate volatility or smaller currencies are more likely to have capital controls.

V. Spreads and Quoting Frequencies

In section II we discussed the different approaches taken to measure forex transaction volume or market activity. One major advantage of taking Reuters FXFX quoting (tick) frequency as a proxy rather than actually reported trading volumes is that the former is available at time intervals much closer to the time horizons of dealers in the market. Bollerslev and Domowitz (1993) conclude their paper by stressing "the potential importance of extending existing literature to replace volume by quote generation activity in order to explain the theoretical link between market activity and the bid-ask spread". However, most recently Goodhart et al. (1996) have thrown some doubt on this claim by comparing half-hourly Reuters FXFX *quoted* data with half-hourly Reuters D2000-2 quoting and *transaction* data for dollar/mark over one day. They discover that D2000-2 *quotes* are a good predictor for D2000-2 *transaction frequency*. However, *FXFX* quotes are a poor predictor for *D2000-2* quoting frequency.

The purpose of this section is twofold. First, we want to show how monthly Reuters FXFX quoting frequency performs as a predictor of realized global trading volumes. Second, we want to test it as a different measure of *expected* market activity in long-run spread estimations. The underlying conjecture is that, although tick frequency might not perform very well for ultra high-frequency (intra-day) estimations, such as Bollerslev and Domowitz (1993) and Goodhart et al. (1996), it might still be a useful measure for longer time horizons. For example, in Hartmann (1996) *unpredictable* quoting frequency has been successfully used as an instrumental variable for unpredictable JPY/USD trading volume.

Turning to the first question, estimations are run with logs of (monthly) volumes as the dependent variable and a constant, logs of monthly Reuters tick frequencies, as well as logs of volatilities as the regressors. We had to remove 3 observations from our original data set, which showed extremely large residuals in preliminary estimations. As one can see from table 1, the tick frequency for PTE/USD and SEK/USD in April 1992 were clearly erratic. In the former case one bank was quoting steadily in very short time intervals, in the latter case two banks seem to have entered a 'quoting war'. In both cases quoting was obviously unrelated to trading activity. Similar problems appear for MYR/USD in April 1989. 52 observations remain, 21 for April 1989 and 31 for April 1992.

Table 4 summarizes the results. By comparing the second and the third column one can see that separated cross-sectional estimations for both periods lead to very different parameter estimates. The ticks parameter in 1992 is twice as large as that in 1989 and the volatility parameter is insignificantly different from zero in 1989 and significantly negative in 1992. In order to get a more precise picture of the form of the structural instability between both dates we again realize F-tests on parameter changes (table 5). Those lead to the conclusion that there is a clear change in slopes, but taken this as given the test of equal intercepts in both periods cannot reject. Because of the instable slopes there is no point in applying the panel data techniques discussed in the two previous sections to the present problem. A pooled regression with time-variant slope dummies for tick frequency and volatility is reported in the last column of table 4 and confirms the results found for the separated estimations.

Table 4: Relationship between volumes, ticks, and volatilities				
Variable/statistic/test		1989 (OLS)	1992 (OLS)	Total (OLS with slope dummies)
Intercept (t, t_{White})		6.314 (3.98, 6.63)	3.688 (2.64, 3.85)	4.913 (4.66,5.51)
Tick frequency	1989 (t, t_{White})	0.536 (4.17, 3.35)		0.575 (4.77, 3.50)
	1992 (t, t_{White})		1.212 (6.96, 10.42)	1.105 (7.02, 9,78)
Volatility	1989 (t, t_{White})	-0.053 (-0.18, -0.25)		0.144 (0.59, 0.52)
	1992 (t, t_{White})		-0.727 (-3.34, -4.78)	-0.780 (-3.54, -4.31)
adjusted R^2		0.46	0.61	0.54
Standard error of regression		1.26	1.19	1.22
White test (p-value)		4.30 (0.51)	4.24 (0.52)	8.70 (0.56)
Jarque-Bera normality test (p-value)		0.18 (0.91)	0.43 (0.81)	0.44 (0.80)

Table 5: Analysis of covariance for volume-ticks estimation				
Test	Chow test (p-value)	Test on joint change in slope parameters, given constant intercept (p-value)	Test on change in ticks parameter, const. intercept and changing volatility parameter (p-value)	Test on change in vo parameter, const. int and changing ticks p (p-value)
	$F_{3,46}$=3.607 (0.02)	$F_{2,47}$=4.573 (0.02)	$F_{1,47}$=8.278 (0.01)	$F_{1,47}$=9.060 (0.00)

We underline that the inapplicability of panel techniques to the present estimations prevents us from accounting for possible individual effects between currency pairs. Keeping this qualification in mind the regressions reported in table 4 suggest that Reuters FXFX tick frequencies have some predictory power for monthly global transaction volumes. Interestingly, the correlation of the number of ticks with transaction volume is stronger (in a numerical as well as in a statistical sense) in 1992 than in 1989. However, the structural instability might indicate that even in more frequent time-series (for single currency pairs) the potential change in the relationship between proxy and actual variable can cause problems, in particular when the time-series covers several years.

There are a number of reasons which could explain the change in the volume-ticks relationship found. One is the potential for changes in Reuters' policy, for example in response to competitive pressure from other financial information services.[14] Another is the sensitivity of quoting frequency to particular events in certain bilateral markets during certain periods. (Notice the explosion of tick frequencies for PTE/USD and SKE/USD in April 1992.) Both points require adequate adjustments in the econometric method or in the data set in order to improve the quality of tick frequency as a proxy for trading activity.

In a final step we want to test tick frequency as an alternative measure of trading activity in our spread estimations. Table 6 shows the results on the volume parameter for the different specifications chosen. (All the other aspects of the spread estimation were unchanged from that in table 2 and are not reported since results hardly differed.) As could be expected, increases in expected tick frequency reduces forex spread, as volume does. Plain FGLS estimations with ticks as the explanatory variable look statistically more significant. Since the frequency of quote updating is probably endogenous, we also instrument ticks with trading volume. This does not change the size of the parameter estimate, but reduces the t statistics. The instrumental variable estimates are very similar to the original FGLS estimates with volumes reported in table 2.

[14] Davé (1993) points to differences in market penetration by Reuters in different geographical areas.

Table 6: Spread estimations using tick frequency (excluding outliers)	
Specification	Volume parameter β_1 (t, t_{White})
Ticks as explanatory variable and no instrument	
One-way random effects model (FGLS)	-0.039 (-1.99, -2.85)
Two-way random effects model (FGLS)	-0.041 (-2.22, -3.63)
Ticks as explanatory variable and volume as instrument	
One-way random effects model (IV)	-0.042 (-1.62, -1.74)
Volume as explanatory variable and ticks as instrument	
One-way random effects model (IV)	-0.067 (-1.48, -2.36)

We also performed another covariance analysis to test for the stability of parameters over time in the specifications with ticks (bottom rows of table 3). Again there is remarkable structural stability over time. This may surprise to some extent, since the volume-ticks relation has been found to be unstable before. In sum, tick frequency performs reasonably well as a measure of trading activity in our spread estimations. It leads to the same qualitative result, saying that - in the long run - more active markets will have lower transaction costs than less active markets. The contradicting or negative results of Bollerslev and Domowitz (1993) and Goodhart et al. (1996) may therefore be due to the rapidly deteriorating information content of tick frequency when shortening intra-day time horizons below certain levels.

VI. Conclusions

In this paper we first discussed the issue of the availability of data to estimate the effect of trading activity on bid-ask spreads in the spot foreign exchange market. Then we undertook two attempts to realize such an estimation using (daily averages of) monthly volume data as sampled by central banks in the most important forex centers and monthly Reuters quoting frequency as measures of trading activity.

The most important results of the paper are that first, available information has not the adequate form to test many existing micro-structural models of forex dealer spreads directly. Second, panel estimations with aggregated data exhibit strong individual effects and weaker time effects. (The individual effects endure after adjustments for exchange regimes or capital controls.) As there is no evidence found that these effects are correlated with the observed explanatory variables a random effects specification to be estimated with FGLS turns out to be the appropriate econometric approach. However, a two-way model achieves only slightly better results than a one-way model (ignoring time effects). Third, while residual variances are higher for 1992 than for 1989, tests on parameter stability clearly suggest that the structural relationship between spreads, volumes, and volatilities did not change over time. This suggests that the evolution of international financial markets (liberalisation, broadening) did not change the basic structural relationship between spreads, volumes, and volatilities but made it a less precise description of reality.

Fourth, the impact of trading activity (measured by turnover or ticks) on Reuters spreads is found to be negative but at a relatively low statistical reliability. Since monthly trading will rather reflect the aggregated predictable parts of trading activities over shorter time horizons, these panel results are complementary to recent time-series analyses showing a significant negative impact of daily predictable volume on average daily spreads (Hartmann, 1996). The impact of exchange rate volatilities on spreads is positive and statistically strong. Finally, on a cross-sectional basis Reuters tick frequency has some predictory power for monthly trading volumes in the forex spot market, but the structural relationship between volumes and ticks is not stable over time.

One interpretation is that we do not find evidence that the growth of the global forex market (BIS 1993) can have adverse effects on the efficiency of this market in the long run. Another interpretation stresses the compatibility of these results with the hypothesis that increased trading makes a currency more attractive as a foreign exchange vehicle, although these economies of scale in vehicle use come out stronger in a very long time series than in the short panel available here.

Two implications for future research come out of the paper. One, real transactions data should be collected from forex dealers and brokers - on a much larger scale than those of Lyons (1995) - such that forex spread theories could be estimated more directly. Two, further tests on the quality of Reuters ticks as a proxy for transaction volume should be realized, in particular at higher frequencies (daily, hourly, etc.). With Reuter's recent switch from FXFX to RICS, which cover more banks and quotes, data quality will probably be enhanced as well. As long as satisfying transaction data are not available using tick frequency might be the most promising approach to analyze the short-run relationship between spreads and market activity.

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Appendix: Panel Techniques[15]

Let i=1,...,N denote the index for currency pairs (N=33), t the time index (t=1,...,T; T=2) and M=55 the total number of observations. y_{it} is the log relative spread for currency pair i at time t, and y is the [55×1] vector of all spread observations. X is the [55×3] matrix of all observations of the independent variables in logs, including a vector of 1's in the first column. The very first step is the OLS estimation of the *pooled* (or *total*) model.

Install Equation Editor and double-click here to view equation. **(A.1)**

β_p is the [3×1] parameter vector and η_p the (non-spherical, normal) error vector. However, as pointed out in section III, individual or time effects may cause spherical disturbances or even parameter biases. We first consider the case of individual effects alone (because it already gives most of the relevant intuition) and then treat the less elegant two-way model more briefly.

1. One-way error components model

For the estimation of a panel specification with *individual effects* alone, i.e. differences in the intercepts between currency pairs which are constant over time, it is useful to proceed in steps from the (one-way) between model through the (one-way) within model to the (one-way) random effects model. In these cases we organise the data such that, going from the top to the bottom, index i runs 'slowly' and index t runs 'fast'. (This will be the other way round in the two-way model below.) The *between* model refers to an OLS regression of a cross-section where every observation is an arithmetic average over both periods.

Install Equation Editor and double-click here to view equation. **(A.2)**

where

Install Equation Editor and double-click here to view equation. **(A.3)**

[15] For an excellent synthesis of most of the following and other panel techniques, see Baltagi (1995).

σ can be useful for the tests and estimations explained further below.

In contrast, in the *within* model each variable is *centered* around the mean over both periods.

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where

Install Equation Editor and double-click here to view equation. **(A.5)**

(Currency pairs with a single observation in time are dropped here.) With only two periods this is equivalent to taking the differences between both periods. The point in doing this is to *remove* any individual effects. In fact, the within model also amounts to the same as introducing a dummy variable for every currency pair, known as *fixed effects* model (e.g. Greene, 1993). Again, estimations are done with OLS.

However, simply 'filtering out' the possible individual effects comes at a high cost in terms of lost degrees of freedom. Therefore, a further step is to estimate a *random effects* model allowing for unobservable individual effects (μ_i) in the stochastic error term.

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where

σ=var(μ_i) ∀i and σ=var(ε_{it}) ∀it. Since the μ_i are a source of autocorrelation in the whole error term η_{itR}, (A.6) has to be estimated using feasible generalized least squares (FGLS). This amounts to the OLS estimation of the following weighted model (see e.g. Greene, 1993).

Install Equation Editor and double-click here to view equation. **(A.7)**

where

Install Equation Editor and double-click here to view equation. **(A.8)**

with

Install Equation Editor and double-click here to view equation. **(A.9)**

and T_i the number of observations for currency pair i. θ can be estimated from the within and the total model. In fact, it can be shown from (A.2) and (A.4) that

Install Equation Editor and double-click here to view equation. **(A.10)**

Since our panel is unbalanced, the denominator of (A.10) cannot be estimated without bias from the between residual variance. Instead we estimate σ indirectly from the residual variances of the pooled (A.1) and the within model (A.4) exploiting

Install Equation Editor and double-click here to view equation. **(A.11)**

Taylor (1980) finds the FGLS estimator for the random effects model parameters to be more efficient than

Install Equation Editor and double-click here to view equation. **(A.12)** the OLS estimator in the fixed-effects (within) model, even for moderately sized samples as the one in this paper. The snag in this specification is that if the random effects (μ_i) are correlated with the (observable) explanatory variables, then * will be biased (Mundlak, 1978). Thus, it has to be tested for the absence of such a correlation between errors and regressors.

The preceding step, however, is to test for the *existence of individual effects* Notice that for balanced panels

Install Equation Editor and double-click here to view equation. **(A.13)**

Thus for the subset of data with two observations in time one can exploit

Install Equation Editor and double-click here to view equation. **(A.14)**

Both estimators (A.11) and (A.14) follow a χ^2 distribution with degrees of freedom (df) corresponding to those of the within and the (adjusted) between (B') estimation. Therefore

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is a statistic with which one can test the null of the *absence* of individual effects (σ=0). If this hypothesis is rejected, then it has to be decided between the within and the random effects specification.

We chose on the basis of two types of tests, a 'Mundlak test' and a Hausman (1978) test. To save space we only describe the specification of the 'Mundlak test' here, as depicted by Dormont (1989).[16] Under Mundlak's hypothesis $E(\mu|X)\neq 0$ in the random effects model (A.6), more precisely

Install Equation Editor and double-click here to view equation. **(A.16)**

By inserting (A.16) in (A.10) one can test the null hypothesis of the absence of a correlation bias with a simple F-test for the linear constraint π=0. If the null cannot be reject, then we can have confidence that,

[16]Strictly speaking, this test was not described in Mundlak's (1978) article. However, it follows directly from the specification chosen by him. For an exposition of the Hausman test see e.g. Greene (1993).

apart from possible other misspecifictions, an FGLS estimation of (A.6) is consistent and asymptotically efficient. However, one should keep in mind that our full sample comprises only M=55 observations. Therefore, statistical inference based on asymptotic arguments should be interpreted with caution.

2. Two-way error components model

The general specification of the two-way random effects model was given in (6) in the main text; adding time effects λ_t, which are constant over individuals, to the one-way model. In the case of unbalanced panels, estimating this model is much more involving than the one-way case because of problems in manipulating the weighting matrix which would allow for the simple application of OLS to the weighted data (see Baltagi (1995) for details). Since most of the intuition is already clear from the one-way case, we limit our attention to the analytical expressions for the appropriate FGLS estimator without going through all the steps to derive it. The exact proofs can be found in Wansbeek and Kapteyn (1989). Following their approach the data are now reorganized with index i running 'fast' and index t running 'slowly'.

Let us first define the two-way within specification, which is equivalent to a fixed effects model with dummy variables for both individual and time effects.

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where

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$\Delta_1 \equiv (D_1,D_2)'_{[55x33]}$, where the D's are obtained from I_N (the identity matrix of order 33) by eliminating the rows which correspond to the currency pairs lacking in period t=1,2. While P is an arbitrary weighting matrix, $\Phi \equiv \Delta_2 - \Delta_1(\Delta_1'\Delta_1)^{-1}(\Delta_2'\Delta_1)$ with $\Delta_{2\,[55x2]}$ being block-diagonal containing two vectors of 1's of order 22 and 33 respectively. Multiplying the spread equation by q filters out both individual *and* time effects. Notice that the standard errors of the OLS regression of the weighted model have to be corrected for the loss of N+T-1=34 degrees of freedom.

The two-way random effects FGLS estimator is where Ω is the covariance matrix of the disturbances of (6) in the main text[17]

σ_w can be taken directly from the within estimation, while the variances of individual and time effects are estimated as $\sigma=[q_1-(N+k_1)\sigma_w-T\sigma]/N$ and $\sigma=[q_2-(T+k_2)\sigma_w-(T/N)(q_1-(N+k_1)\sigma_w)]/(N-T^2/N)$. The following definitions are still needed: $e\equiv y-X\beta_w$, $q_1\equiv e'\Delta_1(\Delta_1'\Delta_1)^{-1}\Delta_1 e$, $q_2\equiv e'\Delta_2(\Delta_2'\Delta_2)^{-1}\Delta_2 e$, $k_1\equiv tr\{(X'qX)^{-1}X'\Delta_1(\Delta_1'\Delta_1)^{-1}\Delta_1 X\}$, $k_2\equiv tr\{(X'qX)^{-1}X'\Delta_2(\Delta_2'\Delta_2)^{-1}\Delta_2\}$.

To test for the joint hypothesis of the absence of both individual *and* time effects - H_0: $\sigma=\sigma=0$, H_1: $\sigma>0$ or $\sigma>0$ - we first apply the simple Honda (1985) test with test statistic

$lm1\equiv M[2(tr\{\Delta_1\Delta_1'\}^2-M)]^{-1/2}[(\eta'\Delta_1\Delta_1'\eta/\eta'\eta)-1]$, $lm2\equiv M[2(tr\{\Delta_2\Delta_2'\}^2-M)]^{-1/2}[(\eta'\Delta_2\Delta_2'\eta/\eta'\eta)-1]$. Then we apply the Gourieroux, Holly, Monfort (1982) test with statistic

The ghm statistic follows a mixture of χ^2 distributions with different degrees of freedom. The critical values of this mixed χ^2 for 10, 5 and 1 % significance are 2.95, 4.32 and 7.29.

[17] A simple weighted regression is not undertaken because of problems in finding an analytical solution for Ω