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Volatility

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The Effect of a Transaction Tax on Exchange Rate Volatility

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Abstract

We argue that a transaction tax is likely to amplify, not dampen, volatility in the foreign exchange markets. Our argument stems from the decentralized trading practice and the presumable discrepancy between 'informed' and 'uninformed' traders' valuations. Since informed traders' valuations are likely to be less dispersed, a transaction tax penalizes informed trades disproportionately, leading to increased volatility. Empirical support for this prediction is found by investigating the effect of transaction costs on the volatility of DEM/USD and JPY/USD returns. High-frequency data are used and an increase in transaction costs is found to have a significant positive effect on volatility.

Keywords: Transaction tax; exchange rates; volatility

JEL-codes: F31, F42, G15, G28

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1 Introduction

There is an ongoing debate about whether transaction costs would increase or decrease the volatility in financial markets. Especially since Tobin's (1978) suggestion of a tax on all foreign exchange transactions intended to dampen alleged excess volatility, a number of studies have been published on this issue. The intuitive argument in favor of a general transactions tax is based on the idea that there are 'noise' traders whose trading is not based on information about security values, and hence, their trades may move prices away from securities' intrinsic values, increasing price volatility. The problem with any indiscriminate transaction tax is, however, that it penalizes not only destabilizing speculative trades but also trades which help to anchor the market by providing liquidity and information. Therefore, as pointed out by several authors (e.g. Kupiec, 1996; Song and Zhang, 2005), the net effect of a transaction tax on price volatility will depend on the change of the trader composition arising from the implementation of the tax. In addition to that, a transaction tax may increase volatility by reducing market liquidity and increasing the price impact of individual trades (e.g. Heaton and Lo, 1995).

There seems to be surprisingly little rigorous theoretical and empirical research on the effects of a transaction tax on volatility, and the existing studies predominantly concentrate on the stock market. As such, all the results obtained in this setup do not necessarily carry over to the foreign exchange market, where trading is still distinctively decentralized. The most noteworthy theoretical studies, i.e. Kupiec (1996) and Song and Zhang (2005), provide mixed results that can be used to rationalize both the conventional 'Tobinian' wisdom as well as the contrarian view. In this article, we argue that, when implemented in the foreign exchange market, a transaction tax is likely to reduce 'informed' trades more than it discourages 'uninformed' agents from trading. As a result, volatility is likely to increase due to an adverse trader composition effect. The empirical evidence presented in the paper lends support to this unambiguous relationship.

Our theoretical argument stems from the decentralized trading practice and the presumable discrepancy between 'informed' and 'uninformed' traders' valuations of the asset. In the foreign exchange markets 'uninformed traders' or 'noise traders' can be viewed as agents who constitute the basic demand and supply in the market. A

noise trader prone to sell (buy) has some ‘idiosyncratic’ reasons to have low (high) valuation for the asset. Among the uninformed traders are importers and exporters in need of a specific currency for their goods transactions. They can, to some extent, be assumed to be unconcerned about whether the currency they demand is fundamentally undervalued or overvalued. On the other hand, ‘informed traders’ can act either as buyers or sellers, depending on their opinion about the intrinsic value of the security and—given the uncoordinated and decentralized trading pattern—whether they contact a trading partner who has lower or higher valuation for the asset. Effectively the introduction of a transaction tax means an equal increase in transaction costs, and any transaction cost will obviously preclude trade between agents whose valuations differ less than the magnitude of this cost. Since it seems justifiable to think that informed traders’ valuations are less dispersed than noise traders’ valuations, a transaction cost is likely to hinder trade most among informed agents. Such an adverse trader composition effect will then increase volatility, because traders with extreme valuations conclude transactions most frequently. Moreover, liquidity is reduced as potentially gainful trades remain unrealized and the price impact of successful trades is weighted.

In line with the existing theoretical literature, also most of the previous empirical work on the effect of transaction costs on financial market volatility concerns stock markets. The studies in this literature make use of time series data around one-time regulatory or institutional changes such as the introduction of or an increase in the transaction tax (Umlauf, 1993; Jones and Seguin, 1997), the introduction of decimal quotation (Bessembinder, 2002) or a tick size change (Hau, 2003). Unfortunately, the foreign exchange market has not experienced similar abrupt institutional changes that could be interpreted as natural experiments and that would allow for empirically testing the implications of our theoretical model. Moreover, a potential problem with the studies of this kind is the difficulty of controlling for other factors that may simultaneously affect fundamental volatility. Therefore, we follow Aliber, Chowdhry and Yan (2003) in examining the effects of transaction costs and interpreting the results in terms of a transaction tax.

Using futures data, Aliber et al. (2003) found a positive relationship between transaction costs and volatility in the markets for the British Pound, Japanese Yen and Swiss Franc (against the U.S. Dollar), in accordance with our theoretical model. However, for the German Mark such a result could not be established. Unfortunately,

their empirical analysis suffers from two potential shortcomings that may explain these somewhat contradictory findings. First their monthly data are highly aggregated: both the transaction costs and volatility can vary considerably within a month or even a day (see Figures 1 and 2) which can obscure the results. To this end, we use daily and five-minute data. Second, as pointed out by Werner (2003), it is possible that the positive relationship between the conditional volatility and transaction cost results from changes in fundamental volatility causing changes in transaction costs. To solve this endogeneity problem an independent measure of fundamental volatility is needed, and we argue that the news count variable included in our data set can better be seen as such a variable than the spot market volatility suggested by Aliber et al. (2003). The empirical results with data on the German Mark and Japanese Yen (against the U.S. Dollar) lend support to our theoretical model: following an increase in transaction costs, price dispersion increases, even controlling for fundamental volatility.

The plan of the paper is as follows. Section 2 formalizes our theoretical argument. The data to be used in the empirical analysis is described in Section 3. The empirical model and results are reported in Section 4. Section 5 concludes.

2 Transaction tax and volatility in the foreign exchange markets: a theoretical perspective

Unlike most other financial markets, the foreign exchange market is still dominated by direct trades between banks acting on behalf of their clients and themselves. According to the Bank for International Settlements (1999), only approximately 25–36% of the transactions (depending on the country) took place through electronic order-matching systems (e.g. Electronic Broking Services (EBS) and Reuters Dealing 2000) in 1998, whereas most of the trade occurred over the telephone based on bid and ask offers conveyed to market participants on the screens of companies such as Reuters, Bloomberg and Bridge. While the share of electronic trading has increased over the recent years, the market is still highly decentralized and can be modeled as a random matching system. Moreover, the empirical evidence in the previous literature as well as in this paper almost solely concentrates on the direct market due to lack of available data on the electronic order-matching systems.

To derive the consequences of transaction costs on price dispersion, we consider the following simple setup that captures the main features of the market. Assume a continuum of traders who differ in their 'idiosyncratic' valuation for the security. The 'type' (i.e. the valuation) of a trader is a realization from the set $\Theta = \{i_1, i_2, n_1, n_2, \}$. The types i_1 and i_2 are informed traders whose valuations are based on their estimate about the 'intrinsic value' of the security. The intrinsic value is normalized to unity. The informed trader i_1 (i_2) has a pessimist (an optimist) opinion about the intrinsic value and his valuation is $1 - \varepsilon$ ($1 + \varepsilon$). The parameter ε captures the difference of opinion between traders i_1 and i_2 . The symmetry of the opinions around the intrinsic value can be assumed without loss of generality because the correctness of the estimate does not play any role in the analysis.¹ The types n_1 and n_2 , in turn, are noise traders with more extreme valuations $1 - x$ and $1 + x$ respectively. We assume $\varepsilon < x/3$ which guarantees that the trader n_1 (n_2) always acts as a seller (buyer) in the market and the difference of valuations between traders is the lowest when the types i_1 and i_2 are matched to trade. For simplicity, we assume that traders are uniformly distributed over the set Θ which means that each trader type is equally common in the market². Moreover, since traders meet randomly, each feasible trading opportunity is equally probable in the market.

Assume first that there are no transaction costs. Upon a meeting, the transaction is concluded at a price that splits the rent (the difference of valuations) evenly among the traders. This practice corresponds to the symmetric Nash bargaining solution. If alike traders meet, there will be no rent to be shared and traders separate without completing the transaction. Table 1 summarizes the prices at which trade is conducted in each feasible trading opportunity.

The expected price is given by

$$E[p] = \frac{1}{6} \left[1 - \frac{x + \varepsilon}{2} + 1 - \frac{x - \varepsilon}{2} + 2 \times 1 + 1 + \frac{x - \varepsilon}{2} + 1 + \frac{x + \varepsilon}{2} \right] = 1,$$

which, due to the symmetry of valuations, equals with the intrinsic value of the asset.

¹The symmetry of the informed valuations also captures the plausible idea that the true intrinsic value can be computed as the weighted average of different opinions; i.e. the pieces of information dispersed in the market aggregates to perfect information.

²This assumption can be made without loss of generality because the *ex ante* composition of traders does not affect the qualitative results.

Table 1: Prices at feasible trading opportunities.

	n_1	i_1	i_2	n_2
n_1	-	$1 - \frac{x+\varepsilon}{2}$	$1 - \frac{x-\varepsilon}{2}$	1
i_1		-	1	$1 + \frac{x-\varepsilon}{2}$
i_2			-	$1 + \frac{x+\varepsilon}{2}$
n_2				-

The variance of the market price obtains

$$\text{Var}[p] = \frac{x^2 + \varepsilon^2}{6}.$$

Assume now that there is a transaction cost C levied on each successful transaction and that this burden is evenly distributed between the traders. The cost C can be interpreted to represent the effects of a transaction tax. Moreover, assume $2\varepsilon < C \leq x - \varepsilon$, so that trade is no longer gainful between the informed agents i_1 and i_2 but is still feasible between the noise traders and between a noise trader and an informed trader. Hence, the trading opportunities between i_1 and i_2 at price $p = 1$ are omitted and other realizations are weighted in the price distribution. The expected price is still $E[p \mid C] = 1$ but the variance yields

$$\text{Var}[p \mid C] = \frac{x^2 + \varepsilon^2}{5} > \text{Var}[p].$$

Thus, as the transaction cost limits the trading opportunities between informed traders, the relative frequency of trades involving at least one noise trader is increased, which then leads to greater price dispersion, too.

In the above analysis we compared, for simplicity, the situations with and without a transaction cost. The fact that in practice there is virtually always a nonzero transaction cost does not change the qualitative conclusions. It is obvious that they generalize to the increase in the existing transaction costs.

3 Measuring Volatility and Transaction Costs in the Foreign Exchange Market

In this section we describe the construction of the data to be used in the empirical analysis. As a starting point we have the high-frequency data set HFDF93 compiled by Olsen and Associates, consisting of the Deutsche Mark–Dollar (DM/\$) and Yen–Dollar (Yen/\$) exchange rate quotations from October 1, 1992 until September 30, 1993. These data have been examined before by Andersen and Bollerslev (1998a, b), *inter alia*. Although the sample period is relatively short, its length is comparable to previous empirical studies employing high-frequency financial data. As the markets are highly liquid, the amount of high-quality data is still more than sufficient for serious statistical analysis. Furthermore, one advantage of this particular data set is that it includes the flow of money-market headline news on the Reuters AAMM screen that can be used to extract the fundamental market volatility to be used as a control variable as discussed below.

Following Andersen and Bollerslev (1998b), we compute a measure of the daily *realized* variance by summing squared five-minute returns over each trading day. The use of five-minute returns is a compromise between the theoretical considerations recommending sampling at very high frequencies and the desire to avoid contamination by microstructure effects (see e.g. Andersen et al., 2001). The returns are computed as percentage differences of the averages of the logarithmic bid and ask prices closest to the end of each five-minute interval. Because of the very small trading activity over the weekends, returns from Friday 21:00 Greenwich Mean Time (GMT) through Sunday 21:00 GMT were excluded, yielding a total of 260 observations of daily realized variance for each exchange rate series. For details of the data set and construction of the series, see Andersen and Bollerslev (1998a).

The measure of the transaction cost is based on the same five-minute bid and ask quotes as the realized variance. Specifically, the daily averages of the following proportional cost measure, C , are used (Aliber et al., 2003),

$$C = \frac{S_a - S_b}{S_a + S_b} \quad (1)$$

where S_b and S_a denote the bid and ask prices, respectively. The rationale behind

this measure can be seen by letting S denote the price of one unit of currency if the customer faces no transaction costs and noticing that the following equalities must hold for the transaction costs to be equal for buying and selling,

$$\frac{S - S_b}{S} = C$$

and

$$\frac{S_a - S}{S} = C.$$

Solving for C now yields (1). In what follows, the percentage measure $100C$ will be used.

The daily transaction cost measures and the realized variance of DM/\$ and Yen/\$ rates are depicted in Figures 1 and 2, respectively. Visual inspection suggests positive correlation between variance and transaction costs. There are also some clearly exceptional periods. The DM/\$ returns were very volatile at the beginning of the sample period, presumably due to the heavy speculation in the market involving several European currencies in September 1992 (see e.g. Andersen and Bollerslev, 1998a). In the transaction cost series, two periods are conspicuous; the highest spike corresponds to Christmas and the second highest to Easter. The market was very thin around these days and the measure of transaction costs is bound to be inaccurate. In the empirical analysis, we use dummy variables to take these as well as some other holidays into account.

4 Empirical Results

4.1 Daily Regressions

To study the relationship between the volatility and transaction costs of exchange rate returns and to test the implication of a positive volatility effect of an increase in transaction costs, realized variance was regressed on transaction costs and some control variables³. Because the realized variance turned out to be autocorrelated in

³We also estimated specifications with the realized standard deviation as the regressand, but according to diagnostic tests these models were not satisfactory.

both cases, its first lag was included in the models. In addition, dummy variables were introduced to take calendar effects into account. First, a Friday dummy was included in all model specifications because specific events in the sample period often happened to occur on Fridays, as pointed out by Andersen and Bollerslev (1998a). Second, there was very little trading around certain national and international holidays, and hence, a dummy variable for these holidays was constructed. This dummy equals unity on the days listed in the appendix to Andersen and Bollerslev (1998a) and zero otherwise. Finally, we also experimented with separate dummies for the most exceptional periods (Christmas and Easter), but they turned out to be insignificant at conventional levels, and despite their inclusion, the parameter estimates remained more or less intact. Thus, the results of these model specifications are not reported.

The results of the OLS regressions are presented in Table 2. For both the DM/\$ and Yen/\$ returns the results of two different specifications are presented to examine the robustness of the results with respect to the holiday dummy. The parameters are, in general, very accurately estimated. However, the holiday dummy is not significant even at the 10% level in the DM/\$ return model, and thus we prefer specification (2). For the Yen/\$ return equation, in contrast, the holiday dummy is highly significant and specification (1) is preferred. In both of these model specifications, the coefficients of the transaction cost variable are positive and significant at the 5% level. Also the lagged realized variance is significant in each case. According to the diagnostic tests, error autocorrelation or autoregressive conditional heteroskedasticity are not a problem and the RESET test does not provide strong evidence against any of the specifications. Errors seem to be heteroskedastic, and therefore, *t*-statistics based on White's (1980) robust standard errors are reported.⁴

In view of the *t*-statistics in Table 2, there indeed seems to be a positive relationship between transaction costs and volatility for both currencies contrary to the findings of Aliber et al. (2003). Thus, the results lend support to our theoretical model in Section 2. In order to get an idea of the economic significance of this relationship

⁴Because it is likely that the error terms of the equations for the DM/\$ and Yen/\$ returns are correlated, efficiency gains could be obtained by estimating the model as a bivariate system. The differences between such GMM estimates and the OLS estimates presented in Table 2 turned out to be minor albeit with somewhat smaller standard errors. Therefore, these results are not reported.

it is interesting to compute the effect of, say, a 0.01% increase in the transaction costs on variance. To this end, we use the estimates of specification (2) for the DM/\$ returns and specification (1) for the Yen/\$ returns. In the former case, taking the effect of the lagged variance into account, the effect of a 0.01% change in the transaction cost equals $0.01\% \cdot 24.197 / (1 - 0.514) = 0.00498$. This is about 1.16% of the average variance of the DM/\$ returns in the sample (0.427). For the Yen/\$ returns the corresponding figure is 1.21%. Both figures indicate that the effect of even a small increase in transaction costs, or equivalently the introduction of an equal tax, would, on average, increase the volatility of the exchange rate returns considerably. For the Yen/\$ returns this is considerably higher than the approximately 0.25% that Aliber et al. (2003) obtained. Furthermore, our estimates for the Yen/\$ and DM/\$ returns are very close to each other, whereas their estimate for the DM/\$ return is clearly smaller and not even statistically significant at conventional significance levels.

It is, of course, possible that the positive relationship between the conditional volatility and transaction cost results from changes in fundamental volatility causing changes in transaction costs. One way to control this effect is to include an independent measure of fundamental volatility as an explanatory variable in the regression model. Recently, Eddelbüttel and McCurdy (1998) have demonstrated, using this same data set, that the frequency of news is strongly associated with volatility, i.e., exchange rates are more volatile during periods with a lot of economic news. Following Eddelbüttel and McCurdy (1998), we thus use the total number of money-market headlines on the Reuters AAMM screen each day. According to the results (not shown), the news count variable is positive and clearly significant (p-values less than 0.0001) for both exchange rates. The other coefficients and their t-statistics are not much affected; the estimated coefficients of the transaction cost variable increase somewhat. Hence, the conclusions drawn above are supported and the causality indeed seems to run from the transaction costs to conditional volatility.

4.2 Intradaily Regressions

Above we have evidence that increases in transaction costs tend to increase daily conditional volatility. The daily data are, however, still quite aggregated in that transaction costs can vary over each day, and, therefore, it might be interesting to check

whether this relationship also holds at the intradaily level. Even though the dataset contains all the quotes, building an operational model beyond the daily frequency is in practice complicated. As discussed by Andersen and Bollerslev (1998a), inter alia, especially intradaily seasonalities and market microstructure effects have to be taken into account. To this end, we employ the Flexible Fourier form (FFF) regression originally introduced by Gallant (1981, 1982) and applied to exchange rate returns by Andersen and Bollerslev (1997). The model is estimated using five-minute returns to mitigate microstructure effects.⁵

The FFF regression equation takes the form⁶

$$\begin{aligned}
2 \log \frac{|R_{t,n} - \bar{R}|}{\hat{\sigma}_t / \sqrt{N}} &= \mu + \delta_{0,1} \frac{n}{N_1} + \delta_{0,2} \frac{n^2}{N_2} \\
&+ \sum_{p=1}^P \left(\delta_{c,p} \cos \frac{2\pi p}{N} n + \delta_{s,p} \sin \frac{2\pi p}{N} n \right) \\
&+ \beta_1 \text{Cost}_{t,n} + \beta_2 \text{Friday}_t + \beta_3 \text{Holiday}_t + \varepsilon_{t,n}
\end{aligned} \tag{2}$$

where $R_{t,n}$ denotes the return in interval n on day t , \bar{R} the sample mean of the five-minute returns and $\hat{\sigma}_t$ an estimate of the daily volatility factor. N is the number of return intervals in a day ($N = 288$) and $N_1 = (N + 1)/2$ and $N_2 = (N + 1)(N + 2)/6$. The trigonometric functions are supposed to capture the smooth intradaily seasonal patterns in volatility. The dummies Friday and Holiday are defined as in the daily

⁵In particular, as Daniélsson and Payne (2002) have recently pointed out, the bid and ask quotes such as ours, obtained from the interbank Reuters network are *indicative* rather than *firm* in that they are not binding commitments to trade. Hence, at very high frequencies they may not accurately measure tradeable exchange rates. Daniélsson and Payne (2002), however, show that at levels of aggregation of five minutes and above, returns computed from these data are a fairly good proxy for firm returns.

⁶This is based on the following decomposition for the intraday returns:

$$R_{t,n} = E(R_{t,n}) + \frac{\sigma_{t,n} s_{t,n} z_{t,n}}{\sqrt{N}}$$

where $\sigma_{t,n}$ is the daily volatility factor, $s_{t,n}$ the periodic volatility component for the n th intraday interval, $z_{t,n}$ an iid(0,1) innovation and N the number of return intervals in a day. For details, see Andersen and Bollerslev (1997).

regressions above, and Cost is computed using formula (1) with the bid and ask prices closest to the beginning of each five-minute interval. Here we are mostly interested in the effect of the transaction costs on volatility (i.e. the coefficient β_1), while the periodic variation in volatility is considered as a nuisance. However, as the results in Andersen and Bollerslev (1997) suggest, ignoring these periodicities could yield misleading conclusions. In practice equation (2) is estimated in two steps. First, a GARCH(1,1) model is estimated for daily exchange rate returns to obtain the daily volatility factor $\hat{\sigma}_t$. Then this $\hat{\sigma}_t$ is plugged into equation (2) which is estimated by ordinary least squares. Note that here we are only modeling the periodic component of volatility, so the magnitude of the estimated coefficients cannot be directly used to compute the effect of the transaction cost on volatility. However, if β_1 turns out to be significantly positive, it indicates that an increase in transaction costs gives rise to increased volatility.

We estimated model (2) using the 74880 five-minute returns of both exchange rates. As expected, the first-step GARCH(1,1) models estimated using daily data from the beginning of October 1987 through September 29, 1993 (not reported) indicated high persistence in conditional variance. The results for the Fourier Flexible form regressions are presented in Table 3. The error term is highly autocorrelated, and following Andersen and Bollerslev (1998a), Newey–West (1987) standard errors with 289 lags are used to compute the t -statistics. The choice of $P = 4$ turned out to be sufficient, and all the coefficients are very accurately estimated. In particular, the estimate of β_1 , the coefficient of the transaction cost variable, is positive and significant at any reasonable significance level for both currencies. Thus these results lend support to the conclusion drawn from daily regressions that an increase in transaction costs indeed seems to lead to higher volatility in accordance with our theoretical model.

5 Conclusion

In this paper, we propose a theoretical model explaining the effect of an increase in transaction costs on price volatility in the foreign exchange market. As a starting point we have the decentralized structure of this market. In the model the market consists of uninformed and informed traders with the former group including importers and

exporters in need of a specific currency. Many of the uninformed traders can thus be assumed to be relatively unconcerned about the ‘intrinsic value’ of the currency they demand. The ‘informed’ traders’ participation, in turn, is motivated by their opinion about whether the currency is fundamentally undervalued or overvalued. We show in a simple decentralized trading model that a transaction tax will have a positive effect on volatility if the informed traders’ opinions are less dispersed than the noise traders’ valuation. This is because the tax (effectively meaning higher transaction costs) hinders ‘informed’ trades disproportionately, giving rise to an adverse trader composition effect.

In support of the theoretical model, we also provide empirical evidence that increases in transaction costs (which proxy the presumable effect of a transaction tax) lead to increased volatility in the foreign exchange market. This relationship seems to be both statistically and economically significant at the daily level, and intradaily regressions reinforce this conclusion, suggesting that it does not result from aggregation effects. The results are in line with the implications our theoretical model but go somewhat contrary to the recent results of Aliber et al. (2003) who, using highly aggregated monthly data, did not uncover a positive relationship between transaction costs and volatility in DM/Dollar market. For the Yen/Dollar market they did find such a relationship although our calculations seem to indicate a somewhat stronger effect. In addition to the level of aggregation, another difference between their study and ours is that we explicitly take into account the effect of changes in the fundamental volatility, thus controlling for the potential endogeneity problem. These findings are bad news to proponents of the Tobin tax whose main argument has been that a transaction tax on foreign exchange would decrease volatility.

References

- [1] Aliber, R.Z., B. Chowdhry, and S. Yan, 2003, Some evidence that a Tobin tax on foreign exchange transactions may increase volatility, *European Finance Review* 7, 481–510.
- [2] Andersen, T.G., and T. Bollerslev, 1997, Intraday periodicity and volatility persistence in financial markets, *Journal of Empirical Finance* 4, 115–158.

- [3] Andersen, T. G., and T. Bollerslev, 1998a, Deutsche Mark–Dollar volatility: Intraday activity patterns, macroeconomic announcements, and longer run dependencies, *Journal of Finance* 53, 219–265.
- [4] Andersen, T. G., and T. Bollerslev, 1998b, Answering the skeptics: Yes, standard volatility models do provide accurate forecasts, *International Economic Review* 39, 885–905.
- [5] Andersen, T.G., T. Bollerslev, F. X. Diebold, and P. Labys, 2001, The distribution of realized exchange rate volatility, *Journal of the American Statistical Association* 98, 42–55.
- [6] Bank for International Settlements (1999), Central bank survey of foreign exchange and derivatives market activity. Basle.
- [7] Bessembinder, H., 2002, Trade execution costs and market quality after decimalization, *Journal of Financial and Quantitative Analysis* 38, 747–777.
- [8] Daniélsson, J., and R. Payne, 2002, Real trading patterns and prices in spot foreign exchange markets, *Journal of International Money and Finance* 21, 203–222.
- [9] Eddelbüttel, D., and T.H. McCurdy, 1998, The impact of news on foreign exchange rates: Evidence from high frequency data, manuscript, University of Toronto.
- [10] Gallant, A.R., 1981, On the bias in flexible functional forms and an essentially unbiased form: the Fourier flexible form, *Journal of Econometrics* 15, 211–245.
- [11] Gallant, A.R., 1982, Unbiased determination of production technologies, *Journal of Econometrics* 20, 285–323.
- [12] Hau, H., 2003, The role of transaction costs for financial volatility: Evidence from the Paris Bourse, working paper, INSEAD.
- [13] Heaton, J. and A. Lo (1995): "Securities transaction taxes: what would be their effects on financial markets and institutions?," in (S. Hammond, ed.) *Securities Transactions Taxes: False Hopes and Unintended Consequences*; Chicago IL, Catalyst Institute.

- [14] Jones, C.M., and P.J. Seguin, 1997, Transaction costs and price volatility: Evidence from commission deregulation, *American Economic Review* 87, 728–737.
- [15] Kupiec, P. H. (1996): "Noise Traders, Excess Volatility, and a Securities Transactions Tax," *Journal of Financial Services Research*, 10, 115-129.
- [16] Newey, W.K., and K.D. West, 1987, A simple, positive semi-definite, heteroskedasticity and autocorrelation consistent covariance matrix, *Econometrica* 55, 703–708.
- [17] Schwert, G.W., and P.J. Seguin, 1993, Securities transaction taxes: An overview of costs, benefits and unresolved questions, *Financial Analysts Journal* 49, 27–35.
- [18] Song, F.M., and J. Zhang (2005), Securities transaction tax and market volatility. *Economic Journal* 115, 1103–1120.
- [19] Tobin, J., 1978, A proposal for international monetary reform, *Eastern Economic Journal* 4, 153–159.
- [20] Werner, I.M. (2003), Comment on ‘Some evidence that a Tobin tax on foreign exchange transactions may increase volatility’. *European Finance Review* 7, 51–514.
- [21] White, H., 1980, A heteroskedasticity-consistent covariance matrix estimator and a direct test for heteroskedasticity, *Econometrica* 48, 817–838.

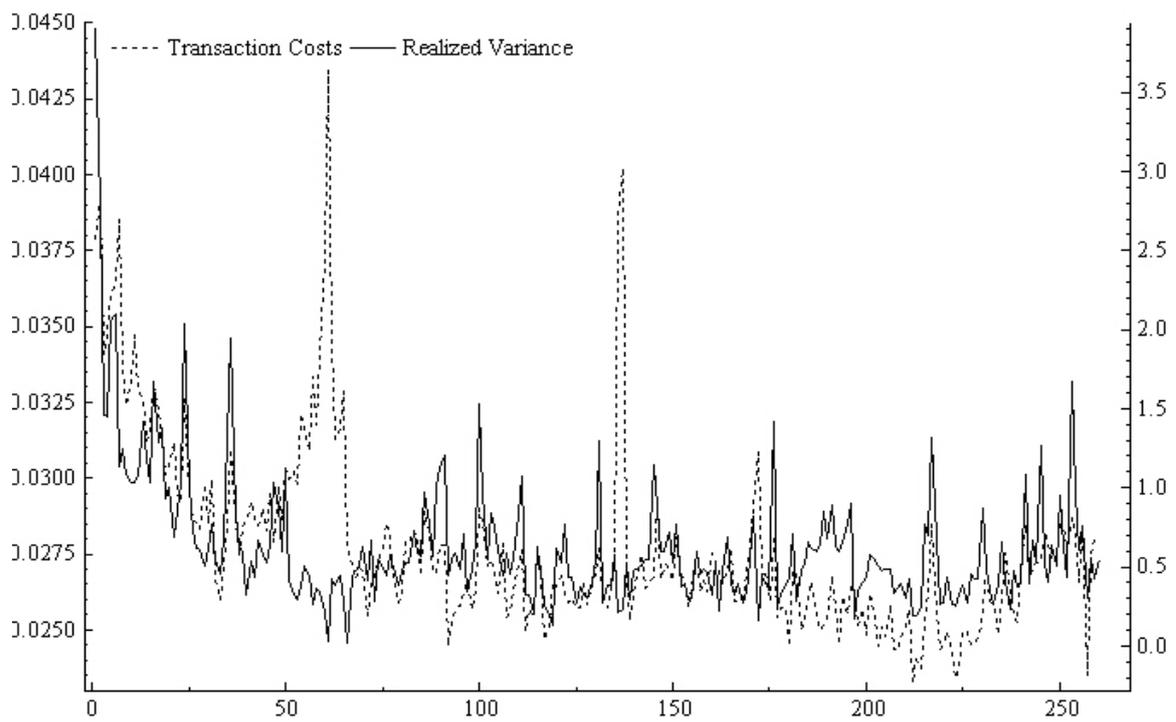


Figure 1: Daily transaction cost (left scale) and realized variance (right scale) for the DM/\$ exchange rate.

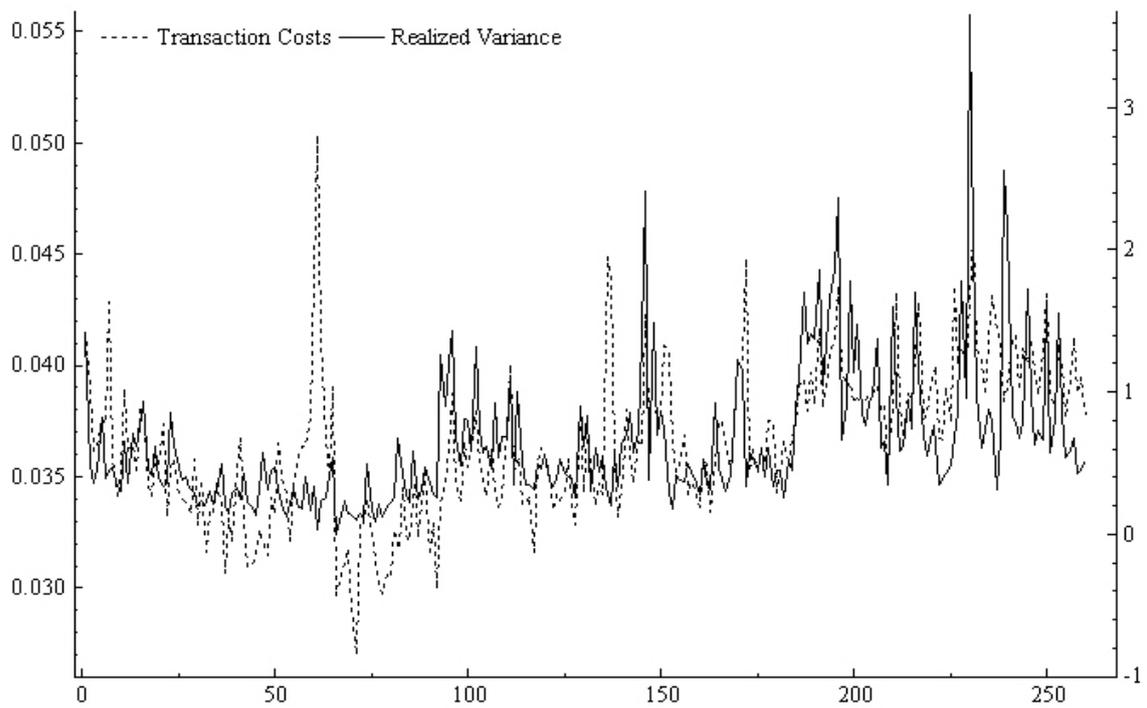


Figure 2: Daily transaction cost (left scale) and realized variance (right scale) for the Yen/\$ exchange rate.

Table 2: OLS regressions of realized variance on transactions costs and control variables.

	DM/\$ Variance		Yen/\$ Variance	
	(1)	(2)	(1)	(2)
Intercept	-0.413 (-1.548)	-0.341 (-1.286)	-1.412 (-3.463)	-1.312 (-3.106)
Transaction Cost	27.506 (2.564)	24.197 (2.294)	51.898 (4.213)	47.825 (3.756)
Lagged Variance	0.499 (7.726)	0.514 (8.179)	0.351 (4.937)	0.388 (5.259)
Friday	-0.187 (-4.397)	-0.184 (-4.344)	-0.218 (-4.050)	-0.209 (-3.910)
Holiday	-0.065 (-1.137)		-0.169 (-2.689)	
R^2	0.480	0.477	0.431	0.417
log likelihood	-29.294	-30.029	-107.394	-110.402
AR(5) ^a	0.076	0.156	0.336	0.214
ARCH(1) ^b	0.108	0.101	0.553	0.540
Heteroskedasticity ^c	<0.001	<0.001	<0.001	<0.001
RESET test	0.097	0.183	0.179	0.592

The figures in parentheses are t-values based on White's (1980) robust standard errors. For the diagnostic tests marginal significance levels are reported.

^aLM test for fifth-order error autocorrelation.

^bLM test for first-order autoregressive conditional heteroskedasticity.

^cWhite's (1980) test for error heteroskedasticity.

Table 3: Results for the Flexible Fourier form regressions.

	DM/\$ Returns		Yen/\$ Returns	
μ	1.584	(2.512)	-3.166	(-4.375)
$\delta_{0,1}$	-9.152	(-5.013)	-7.838	(-3.804)
$\delta_{0,2}$	3.076	(5.035)	2.635	(3.821)
$\delta_{c,1}$	-2.155	(-5.842)	-1.778	(-4.276)
$\delta_{s,1}$	-0.836	(-13.601)	-0.437	(-6.377)
$\delta_{c,2}$	-0.513	(-5.493)	-0.483	(-4.533)
$\delta_{s,2}$	0.077	(1.969)	0.106	(2.673)
$\delta_{c,3}$	-0.537	(-11.195)	-0.518	(-9.630)
$\delta_{s,3}$	0.467	(15.424)	0.413	(12.589)
$\delta_{c,4}$	-0.169	(-4.874)	-0.243	(-6.271)
$\delta_{s,4}$	-0.330	(-11.934)	-0.345	(-12.080)
β_1	17.064	(3.994)	129.715	(5.954)
β_2	-0.266	(-2.213)	-0.330	(-2.847)
β_3	-0.423	(-1.991)	-1.052	(-3.442)

The figures in parentheses are t-values based on Newey and West's (1987) robust standard errors incorporating 289 lags.