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Arbitrage, Covered Interest Parity and Long-Term Dependence between the US Dollar and the Yen

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Abstract

Using a daily time series from 1983 to 2005 of currency prices in spot and forward USD/Yen markets and matching equivalent maturity short term US and Japanese interest rates, we investigate the sensitivity over the sample period of the difference between actual prices in forward markets to those calculated from short term interest rates. According to a fundamental theorem in financial economics termed covered interest parity (CIP) the actual and estimated prices should be identical once transaction and other costs are accommodated. The paper presents four important findings: First, we find evidence of considerable variation in CIP deviations from equilibrium that tends to be one way and favours those market participants with the ability to borrow US dollars (and subsequently lend yen). Second, these deviations have diminished significantly and by 2000 have been almost eliminated. We attribute this to the effects of electronic trading and pricing systems. Third, regression analysis reveals that interday negative changes in spot exchange rates, positive changes in US interest rates and negative changes in yen interest rates generally affect the deviation from CIP more than changes in interday volatility. Finally, the presence of long-term dependence in the CIP deviations over time is investigated to provide an insight into the equilibrium dynamics. Using a local Hurst exponent – a statistic used in fractal geometry - we find episodes of both positive and negative dependence over the various sample periods, which appear to be linked to episodes of dollar decline/yen appreciation, or vice versa. The presence of negative dependence is consistent with the actions of arbitrageurs successfully maintaining the long-term CIP equilibrium. Given the time varying nature of the deviations from equilibrium the sample period under investigation remains a critical issue when investigating the presence of long-term dependence.

Keywords: Hurst exponent; Efficient market hypothesis; covered interest parity, arbitrage

JEL classification: C22; C32; E31; F31

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Arbitrage, Covered Interest Parity and Long-Term Dependence between the US Dollar and the Yen

1. Introduction

A fundamental theoretical premise in financial economics concerns the assumption that equilibrium exists between spot and forward exchange rates and their underlying interest rate markets. Covered interest parity (CIP) arbitrage ensures that equilibrium prices in forward currency markets are maintained based upon interest rate differentials. The objective of this study is to investigate the sensitivity of the residuals – or "observed divergences" (Cosandier and Lang, 1981) or "deviations" (Balke and Walke, 1998)- between estimated forward rates calculated from interest rate differentials and the actual forward rate that exists simultaneously in the market, to time related dependencies. There is an extensive existing literature that has observed systematic deviations from CIP and attributed this anomaly to institutional features such as taxes, market illiquidity, political risk and transaction costs (see Taylor, 1987 and 1989; Poitras, 1988; Popper, 1993; Crowder, 1995; Strobel, 2001; Peel and Taylor, 2002; amongst others).

As suggested by Poitras (1988) and Taylor (1989) amongst others, we first accommodate many of the institutional factors, which may have affected the results of earlier researchers through the use of contemporaneous Euromarket interest rates and similar credit rated bank prices for both spot and forward exchange rates. This investigation, across a 25-year period, also employs a longer daily dataset than other investigations. The currency market investigated is the spot and forward U.S. dollar to Japanese yen (US$/¥)[3] exchange rate. In April 2004 the Bank for International Settlements (BIS, 2005) reported that the spot US$/ ¥ currency pair accounted for 17% (US$296 billion) of daily turnover in spot and forward markets, second only to the euro at 28% (US$501 billion). This currency pair was chosen, since it has the longest time series of the major currencies; the euro being introduced in 2000, and has the next highest level of liquidity.

The approach of this study is also comprehensive in that three different aspects of the deviations from parity are investigated. First, the time varying nature of the parity deviations is estimated. Importantly, while the deviations vary considerably over the 25-year sample period and peak –as might be expected- during periods of economic turbulence, such as occurred during the Asian Crisis of 1998, they are found to have become almost non-existent in recent years; the likely consequence of electronic trading and pricing systems. Second, regression analysis is employed to determine the relationship between the deviations from parity and macroeconomic phenomena such as interest and exchange rate volatility. In this instance we find that interday negative changes in spot exchange rates, positive changes in US interest rates and negative changes in yen interest rates generally affect the parity deviations more than changes in interday volatility in these assets.

Finally we investigate the equilibrium relationship using the statistical tools from fractal geometry. Rather than focus on a cointegration approach (Vierira, 2003),

[3] This is expressed as 1US$ = *x* Yen, which is the convention for this currency pair in international markets.

which aims to identify the presence of a long-term equilibrium, we investigate the time-varying dependence in the residuals using the classical rescaled adjusted range technique of Hurst (1951) after accommodating the effects of short-term autoregression and conditional volatility. This approach differs from cointegration techniques by providing an insight into the direction of the equilibrium reverting process that underpins CIP, since a negatively dependent series is one that reverts to its long-term mean after a shock away from the mean, whereas a positively dependent series is one that progressively moves away from the long-term mean and equilibrium after each new shock.

The paper is set out as follows. In the next section a review of the arbitrage relationships in international financial markets is briefly undertaken, the data and the method for calculating dependence in the residuals is described, and then the results for the econometric analysis and Hurst specifications are presented. The final section allows for some concluding remarks.

2. Arbitrage Relationships in International Finance

There are a number of parity relationships concerning the interaction between spot and forward exchange rates and the underlying forward margins and interest rates, commonly tested in financial economics. These relationships can be simplified into three main groups: (a) equilibrium relationships that are immediately arbitrageable; (b) those relationships, which theory proposes will reach equilibrium over a longer-term horizon; and (c) those which provide a forecast for future asset values.

While the focus of this paper is on covered interest parity, an arbitrageable relationship that can be exploited immediately, to place this literature in perspective it is worthwhile briefly reviewing some key studies in the other two areas mentioned. Addressing the later relationship first, forward exchange rates are said to generally provide unbiased estimates of future spot rates. Previous work in this area is illustrated by the investigations of Phillips and McFarland (1997), Smoluk et al. (1998) and Newbold et al. (1998), which show that while the unbiasedness hypothesis is often rejected, the forward rate remains an important predictor of the future spot rate. These may also be better than survey forecasts, which Elliott and Ito (1999) find were worse than random walk predictions in terms of mean square forecast errors.

Tauchen (2001) points out the pure expectations theory of unbiased forward exchange rates predicts that the slope coefficient in a regression of the change in the spot rate on the difference between the current forward and spot rates should equal unity. In this context his study finds that the sampling distribution of the regression estimator of this coefficient is upward-biased relative to unity and strongly skewed to the right. Nonetheless Wu and Chen (1998) attribute the diversity in many empirical findings to forecasting errors that are sufficiently large, such that a ‘correct’ or an ‘unreasonable’ estimate of the mean value of the exchange rate change renders almost identical results.

The related analysis of the uncovered interest parity relationship –where the interest differential between two countries should equal the expected exchange rate change – may be illustrated by the recent study by Chaboud and Wright’s (2005). In this case the authors note that uncovered interest rate parity is consistently rejected by the data with the higher yielding currency usually appreciating, rather than depreciating, as theory would suggest, over time. Nonetheless, the authors find that over very short time horizons uncovered interest parity hypothesis is in fact supported.

The second relation suggests that in the longer term spot exchange rate should reflect inflation and the relative exchangeability of assets between countries; Turtle and Abeysekera (1996) and Moore and Roche (2001) provide a comprehensive discussion of these relationships. The study by El-Gamal and Ryu (2006) also provides general support for the long-term Purchasing Power Parity (PPP) hypothesis. More recently, with high frequency datasets readily available, researchers, notably in econophysics, have found that spot markets contain some predictability in their market dynamics. For example, Ohira et al.'s (2002) analysis of high frequency exchange data, finds that the return dynamics of the spot US$/¥ is not completely random and that a probabilistic structure exists.

The focus of this paper is on one of two key relationships in cash-based foreign exchange markets that are immediately arbitrageable: covered interest parity and triangular arbitrage. Trading in these markets have shifted over the last two decades from broker driven over-the-counter markets to mostly electronic trading via dealing systems such as Reuters D2000 (see Payne, 2003 for a discussion) and the EBS trading platform. Importantly for this study, the takeup of these systems by market participants has increased in recent years so that they now dominate trading practice. Prices from these trading platforms now also feed into other pricing systems such as those for calculating FRAs and interest rate swaps.

Triangular arbitrage takes advantage of the three possible exchange rate pairs (cross-rates) that exist between three currencies. The triangular arbitrage transaction ensures the product of the three exchange rates is a certain value. In their study, employing high-frequency foreign exchange data, Aiba et al. (2002) demonstrate that triangular arbitrage opportunities exist in the foreign exchange market such that the rate product μ fluctuates around a value m. This last point is particularly important since this suggests the three spot currency pairs are also correlated to one another. The later study by Aiba and Hatano (2004) argued that selective trading generates correlation even without the presence of a triangular arbitrage transaction. Aiba et al. (2003) also show that the negative auto-correlation evident over the short term in high frequency data sets maybe due to the effect of triangular arbitrage.

With respect to covered interest parity and following the study by Frenkel and Levich (1981) and Popper (1993), begin by expressing the relation between the spot (e_s) and forward (e_f) exchange rates and the underlying interest rates over a specific maturity (m) as

$$(1 + i_m) = e_{sm}/e_{fm}(1+i_m^*) \qquad (1)$$

where i_m and i_m^* are the respective domestic and foreign interest rates on securities with the same maturity as the forward rate. More precisely the spot exchange rate (t_s) requires cash settlement two working days from its trade date (t_0) and is expressed as one unit of domestic, or home, currency in terms of a specific amount of foreign currency (e_s). The forward exchange rate (e_f) is expressed the same way as the spot rate, is also observable at t_0, but instead requires cash settlement at a future date t_f. The maturity of these contracts (the number of days between t_0 and t_f) is commonly expressed by market convention in weeks and months from spot. In addition, two sets of interest rates i_m for the home currency rate and i_m* for the foreign interest rate, represent either the cost, or investment return, from either borrowing, or lending, for the *m* period. Importantly, this equation assumes that financial market participants have equal access to these capital and foreign exchange prices, although the effect of segmentation, as discussed by Blenman (1991) is to establish the presence of a forward price band around an equilibrium price.

Covered interest parity (CIP) arbitrage ensures that equilibrium prices in forward currency markets are maintained based upon interest rate differentials. That is, by rearranging (1) the equilibrium forward premium (or discount) on the spot foreign exchange rates is therefore the interest rate differential

$$(i_m^* - i_m)/ (1+i_m^*) = (e_{fm} - e_{sm}) / e_{sm} \tag{2}$$

In practice, this mathematically simple calculation requires consideration of the different money market bases (either a 360, 365 or actual number of days –to accommodate leap years) that exists by convention in different financial markets. Also, arbitrage requires undertaking actual cash flow in all currency positions, which may add to the transaction costs and impose boundaries around the equilibrium price. To some extent these costs may be avoided or reduced using derivatives such as options (Ghosh and Ghosh, 2005).

The objective of this study is to investigate the sensitivity of the deviations from equilibrium (δ) between the estimated forward rate based upon the interest rate differentials (e_f*) and the actual forward rate e_f that is quoted at t_0 in the foreign exchange market to time dependence and related factors. In this way we do not specifically consider the transaction band associated with two-way quotes (due to the bid-ask spread) in foreign exchange markets and the associated algebra (see Balke and Wohar, 1999). Specifically, from equation (1) the estimated forward rate based upon interest rate differentials should be equivalent to the observable forward rate

$$e_{fm}^* = e_{sm} (1+i_m^*) / 1(1 + i_m) \equiv e_{fm} \tag{3}$$

with the deviation from equilibrium (δ_m) being simply the difference between the actual and estimated forward rate for a specific maturity

$$e_{fm} - e_{fm}^* = \delta_m \tag{4}$$

What is of interest in this paper is the behaviour over time of the residuals δ_m, which ideally should follow a random and i.i.d. $N(0, \sigma^2)$ process since CIP arbitrage if maintained over time should cause the deviations to revert to an equilibrium around zero (or close to zero if there is a trading band). Empirical evidence however does not support such a claim. In an earlier study Cosandier and Lang (1981) find the distribution of the arbitrage margins to be non-normal, while Taylor (1987) and Blenman (1991) find a no-arbitrage band within which deviations are random, outside of which deviations revert to the edge of the band. A number of authors observe that the degree of deviation over time is both time varying and also a function of the maturity of the arbitrage investigated. Thus, while Taylor (1999) finds evidence of a maturity effect in shorter dated bill markets, Fletcher and Taylor (1996) and Popper (1993) find evidence of persistent deviations in longer maturities. Later, Poitras (1988) when investigating the CIP relationship on the US$ - Canadian dollar noted that the presence of the arbitrage boundaries (the likely consequence of transaction costs and market segmentation) would affect the residual distribution and recommended the use of an autoregressive model (AR) to correct for residual persistence and any permanent components. These issues are developed further in the next section.

3. Data and Method

3.1 Data

London interbank spot and forward foreign exchange midrates on the US$/¥ and Euromarket yen and US dollar interest rates with 1 month, 3 month, 6 month and 1 year maturities were chosen to investigate the CIP relationship on the US$/¥. All prices were at the daily close of trading. Originally series from the 1 January 1983 to the 7th April 2005 were downloaded from Datastream. Due to some incomplete series for the forward and money market rates, the starting date of the series was made the 11 October 1983, for a total of 5603 daily observations. Our understanding is that this is the longest sample period yet tested for CIP. Implied forward rates, for a specific maturity (m) based upon US ($) and yen (¥) interest rates were calculated based upon Equations (3, 4) and estimated consistent with Taylor's (1989) discussion:

$$\delta_m = e_{fm} - e_{sm}(1+i^{¥}_m) / 1(1 + i^{\$}_m) \qquad (5)$$

(Insert Figure 1 about here)

Figure 1 illustrates the process of actually executing a CIP arbitrage. From Figure 1, begin by either buying or selling US$ spot against yen (top left and right hand corners of the Figure). This results in either a positive or negative spot cash flow in US$ and the reverse cash flow in yen, which then must either be invested or borrowed. In practice, the bid-ask spread, representing the market offer (for you to borrow) and bid (for you to lend) is commonly 1/8th of a percent on Euromarket deposits. Initially, we simply use midrates – and these are what are reported in the subsequent Tables – although simply adding or subtracting 1/16th of a percent to the midrate can recreate the underlying bid and offer rates. In the case of the spot and forward exchange rates, while midrates are also used for the reported calculations, the 5 basis point bid-ask spread typically required by traders can also be accommodated by adding or subtracting a 2.5 basis point spread from the midrate.

The resulting cash flows in spot markets now sum to zero, with the remaining spot cash flows now occurring at the future date being the maturity of the loan or the borrowing. The implied forward rate can easily be derived by dividing the future yen cash flow with the future US$ cash flow. This implied rate might then be compared with the actual forward market rate (bottom left and right hand corners of the Figure). If markets are in perfect equilibrium (and no transaction costs) then the difference (δ_m) should be zero. If δ_m is positive, that is $e_{fm} > e_{fm}$ * then an arbitrage can be executed which requires selling e_{fm} and buying e_{fm} *. The long e_{fm} * position can be created by buying US dollars spot against yen, lending US$ and then borrowing yen. The opposite is also true; if δ_m is negative, that is $e_{fm} < e_{fm}$ * then an arbitrage can be executed, which requires buying e_{fm} and selling e_{fm} *. The short e_{fm} * position can be created by selling

US dollars spot against yen, borrowing US$ and then lending yen. Segmentation in interest rate markets due to credit constraints might prevent access to one particular market. For example, Poitras (1988) noted that in the US-Canadian dollar market limited access to US dollar borrowing ensured that CIP arbitrage tended to be one way, favouring those with the ability to borrow in US interest rate markets.

3.2 Regression Analysis of Deviations from Equilibrium

The underlying parity relationships suggest that deviations from equilibrium may be due to mispricing in interest rate and exchange markets, the presence of a transaction band, or volatility in underlying asset markets. To investigate these possible effects we investigate the following regression model with a GARCH(1,1) specification using methods described by Bollerslev and Wooldridge (1992) to investigate the time varying volatility structure of the return series.. To stabilise δ_m first differences ($\Delta\delta_{mt} = \delta_{mt} - \delta_{mt-1}$) are employed. Logarithmic values are not taken due to negative interest rates on some yen securities in recent years. Thus for a specific CIP maturity (m) at time t (both t and m are dropped to simplify the expression), where

$$\Delta\delta = c + B_1\Delta e_s + B_2\Delta i^{\$} + B_3\Delta i^{¥} + B_4\Delta\sigma(e_s) + B_5\Delta\sigma(i^{\$}) + B_6\Delta\sigma(i^{¥}) + B_7\Delta\delta_{t,t-1} + B_8\Delta\delta_{t-1,t-2} + B_9\Delta\delta_{t-2,t-3} + \varepsilon$$

$$\sigma^2 = \alpha + \beta\,\varepsilon^2_{t-1} + \gamma\sigma^2_{t-1} \tag{6}$$

c	= the regression constant
$B_1\Delta e_s$	= the interday change in the spot rate with this rate expressed in terms of 1US$
$B_2\Delta i^{\$}$	= the interday change in the US$ Euromarket interest rate
$B_3\Delta i^{¥}$	= the interday change in the US$ Euromarket interest rate
$B_4\Delta\sigma(e_s)$	= the interday change in 1 month (22-day) lagged volatility in the spot rate expressed as a percent
$B_5\Delta\sigma(i^{\$})$	= the interday change in 1 month (22-day) lagged volatility in the US$ Euromarket interest rate expressed as a percent
$B_6\Delta\sigma(i^{¥})$	= the interday change in 1 month (22-day) lagged volatility in the Yen Euromarket interest rate expressed as a percent
$B_7\Delta\delta_{t,t-1}$	= an autoregressive term (of the dependent variable) at lag one
$B_8\Delta\delta_{t-1,t-2}$	= an autoregressive term (of the dependent variable) at lag two
$B_9\Delta\delta_{t-2,t-3}$	= an autoregressive term (of the dependent variable) at lag three
ε	= the regression residual
α	= the long-term average of the conditional volatility structure
$\beta\,\varepsilon^2_{t-1}$	= the ARCH term, which represents the significance of volatility observed in the previous period;
$\gamma\sigma^2_{t-1}$	= the GARCH term, which represents the forecast variance from last period.

3.3 Estimating Long-Term Dependence

Many researchers employ cointegration tests of covered interest arbitrage to establish whether a long-term equilibrium exists between one or more economic relationships. For example, Abeysekera and Turtle (1995) and Moosa and Bhatti (1996) demonstrate that CIP maintains the cointegration relationship, although rejections are frequent and robust to both subperiod analysis and alternative interest rates series. Nonetheless, although test rejections are statistically significant, economic profits are likely to be small. Vierira (2003) utilises long-term swap rates to establish the level of capital mobility within the EU. Others focus on analysis of the dynamics of prices either side of the transaction cost band. These empirical studies find considerable differences in the dynamics inside and outside the transaction band. For example, Balke and Wohar (1998) investigating CIP in the pound to US dollar, find that while the impulse response functions inside the transaction costs band are nearly symmetric, those outside the bands are asymmetric-suggesting less persistence outside of the transaction costs band than inside the band.

The presence of non-linear dependence in financial markets, which is a departure from the fair game, or martingale, property expected under market efficiency is also a common feature of financial time series. A statistical property of importance and one that is investigated in this paper is the presence of long-term dependence in the time-series (Ambrose, Ancel and Griffiths, 1992). The implications of dependent processes, evident from low and high order autocorrelation structures in the data are of particular concern for the volatility based pricing models (such as option pricing models) typically used in financial markets. Low order correlations, which tend to exhibit hyperbolic decay, may be associated with short-term memory effects, while long-term memory effects have been linked to the presence of fractal structures.

Long-term dependence in time-series data may be measured using statistical techniques based on range analysis where estimates of dependence using the classical rescaled adjusted range technique of Hurst (1951) yield an exponent (H), which under the assumption that the series follows a Gaussian random walk, equals 0.5. The failure to identify long-term dependent effects also lends support to the proposition that the time series conforms to normally distributed standard Brownian motion. Evidence supporting dependence in currency markets is extensive and includes the work of Muller et al (1990), Byers and Peel (2001) and Batten, Ellis and Hogan (2005), amongst many others. The statistical method employed for measuring long-term dependence in this study is based on the classical rescaled adjusted range (Hurst, 1951; Mandelbrot and Wallis, 1969; Lo 1991).

Begin by taking the δ_{mt} deviation from equilibrium for a specific CIP maturity, filtered for the presence of first order regressive (AR1) features leaving a residual ψ at time t

$$\delta_{mt} = \beta(\delta_m)_{t-n} + \psi_t \tag{7}$$

For each ψ_t the classical rescaled adjusted range $(R/\sigma)_n$ is calculated as

$$(R/\sigma)_n = (1/\sigma_n)\left[\underset{1 \le k \le n}{\text{Max}} \sum_{j=1}^{k}(X_j - \overline{X_n}) - \underset{1 \le k \le n}{\text{Min}} \sum_{j=1}^{k}(X_j - \overline{X_n})\right] \tag{8}$$

where $\overline{X_n}$ is the sample mean $(1/n)\Sigma_j X_j$ of a ψ_t nd σ_n is the standard deviation of ψ_t over a particular series n

$$\sigma_n = \left[1/n\sum_{j=1}^{n}(X_j - \overline{X_n})^2\right]^{0.5} \tag{9}$$

In order to capture the time-varying nature of dependence in ψ_t this study employs a local measure of the Hurst exponent (h). Calculated as

$$h_n = \frac{\log(R/\sigma)_n}{\log n} \tag{10}$$

Under the null hypothesis of no long-term dependence, the value of $h_n = 0.5$. For time-series exhibiting positive long-term dependence, the observed value of the exponent $h_n > 0.5$. Time-series containing negative dependence are alternatively characterised by $h_n < 0.5$. Importantly for an equilibrium reverting process such as CIP, the local Hurst exponent should be negative, since a movement back towards the equilibrium should follow a movement away from equilibrium. For positively dependent processes another movement further away from equilibrium will follow the earlier movement away from equilibrium. These movements could be either negative or positive representing CIP arbitrage in favour of US borrowers or Yen borrowers.

Estimates of the local Hurst exponent are calculated for (N – n + 1) times overlapping subseries of length n, with n having a set value. In this case, n is arbitrarily set to either 22 days or 66 days, which is equivalent to a standard one and 2 month period. The procedure in effect creates a time-series of exponent values, the change in whose value can be measured over time. Note that changes in the local Hurst exponent for each ψ_t may be due to either a time-varying range (R_n) or standard deviation (σ_n), or both of these simultaneously. Thus values for the range, standard deviation and local Hurst exponent are separately recorded for the series.

3.4 Time variation in local Hurst statistics

To provide economic meaning to the stream of Hurst statistics generated across the sample period, we conveniently bundle the estimated Hurst exponent (and its components) into monthly and yearly averages. Analysis of variance (ANOVA) is then used to investigate and model the relationship between a response variable and one or more independent variables. The method extends the two-sample *t*-test for the equality of two population means to a more general null hypothesis of comparing the equality of more than two means, versus them not all being equal. That is the null hypothesis of similarity between groups, k where the groups are monthly or yearly intervals is

$$H_0: \quad \mu_1 = \mu_2 = \ldots \mu_k \qquad \forall\, k$$
$$H_1: \quad \mu_1 \neq \mu_2 \neq \ldots \mu_k \qquad \forall\, k \tag{10}$$

The *F*-test *p*-value is employed to indicate the degree of significant of differences among the means.

(Insert Table 1 about here)

4. Results

The results of the analysis are presented in a series of Tables and Figures. The first, Table 1, provides the yearly mean and standard deviation of the CIP deviation from equilibrium over the sample period. A simple F-test demonstrates statistically significant variation in these annual averages, with the underlying series displaying slight skewness and significant kurtosis. Interestingly, the time variation in these CIP residuals appears to face an apparent boundary at zero, a result consistent with Poitras (1988) on the US$/CAD$. This in effect provides a one-way (unidirectional) CIP arbitrage that favours US$ borrowers.

There is also evidence of a maturity effect (Taylor, 1987, 1989) since the average deviation from parity is greatest for the 1-year (-0.10026) and least for the 1-month (-0.02324). Note that this pattern persists across all years and maturities only differing in its scale, with the greatest deviation in 1998-during the Asian crisis when there was significant US$ appreciation and market "turbulence". Interestingly, the current levels (period 2000-2005) of CIP deviations are low historically and may be the consequence of improvements in technology and the wide adoption of electronic trading platforms such as the Reuters D2000.

(Insert Figures 2, 3)

The monthly mean and standard deviation of the CIP deviations reveal a more complex story than suggested by the simple annual analysis. To illustrate these findings, Figures 2 and 3 plot the monthly mean and standard deviation of the deviation from CIP equilibrium over the sample period from 1983-2005 for 1, 3, 6-month and 1-year maturities. The horizontal axis is in years, while the y-axis is the deviation from CIP equilibrium in basis points. Interestingly there is now evidence of bidirectional arbitrage, a finding consistent with Kia (1996) and Balke and Wohar (1999), although in this instance on an annual basis the negative deviations appear to dominate.

Note from these figures that the period 1997 to 2000 –the period of the Asian Crisis- shows the greatest mean deviations from CIP equilibrium for the entire sample period. The disruptive effect on the CIP equilibrium of certain –usually catastrophic - economic events is also consistent with the findings of other researchers such as Balke and Wohar (1999), who pointed out the impact of the withdrawal of the pound from the European Monetary System in September 1992 on the US-Pound CIP relation. It is also clear from this figure that the mean deviations from CIP have visibly diminished in the period after 2000 – and there has also been a visible drop in volatility. This is confirmed with the yearly averages recorded in Table 1- the lowest mean deviations are for the 2000-2004 years.

(Insert Table 2)

Table 2, which shows the deviations from CIP for specific maturities on the USD/Yen for each month in the year from 1983 to 2005, provides some evidence of a calendar effect. In this case the monthly deviations from CIP are least in the Christmas period and greatest in the summer period, with the deviations are least for the 1-year maturities and most for the 1-month maturities.

(Insert Table 3)

Table 3 reports the regression results of the interday changes in CIP residuals across various time periods 1983 to 2005. Beginning with the results from the total sample, the adjusted R^2's with values from 0.23 to 0.40 suggest that overall the mean equation provides a reasonable explanation of variation in the CIP deviations from equilibrium over the sample period. The series all have significant AR(n) terms, which drop away quickly after one day although this suggests that there is persistence in the arbitrage that at least spans 2-3 days. There are also significant GARCH (1,1) effects with the GARCH (1) term generally more significant than the ARCH (1) term. A higher order GARCH term (not reported for the sake of brevity) was often required in some cases to stabilize the volatility equation. The GARCH coefficient α, is generally not economically significant for any of the maturities, whereas the coefficient β is more significant in the shorter maturities (1,3 months), while the GARCH coefficient γ is more significant in the longer 6-month and 1-year maturities.

Considering individual explanatory variables in the regression equation, the results suggest that interday negative changes in spot exchange rates, positive changes in US interest rates and negative changes in yen interest rates generally matter more than changes in interday volatility in these assets. These results are maintained during different sample periods as well. In the case of the volatility measures the importance of changes in the spot exchange and interest rate volatilities differ over the various sample periods, although consistently the volatility in US interest rates was tenfold higher than volatility in yen interest rates. Note that the signs of the volatility coefficients were generally negative. This suggests that a positive increase interday in volatility will lead to an increase in the negative deviation from CIP equilibrium, whereas a decline in volatility (a negative sign) will reduce the negative deviation from equilibrium.

Of relevance for policy makers and central banks is the negative sign on the spot yen variable (associated with yen appreciation). This is consistent with spot market shocks (such as central bank intervention), which favour the appreciation of the yen, disrupting the equilibrium in forward markets. It appears that US dollar spot appreciation does not have the same destabilizing effect. This relationship persists across the subsequent subperiod analysis.

To understand these relationships it is appropriate to recall the earlier Tables (also Figures 2 and 3) that show that deviations from equilibrium tend to be negative; that is the arbitrage tends to be one way favouring those with access to US dollar borrowings– since the implied forward rate tends to be higher than the implied forward rate based on interest rate differentials. Thus, consistently positive interday changes in the US Euromarket interest rate (associated with interday increases in US interest rates) appear to inhibit access to Euromarkets. This may be a timing related effect due to Japanese and US markets occupying different time zones, even though the prices used for this study are all London prices.

(insert Table 4)

The next step in the statistical analysis requires the investigation of long-term dependence in the CIP residuals using the tools developed in fractal analysis. In this case the local Hurst coefficient is calculated after filtering for AR(1) effects. This is easily justified given persistence in the AR(n) terms described earlier. Table 4 reports the annual deviations from CIP over the sample period; in this case the 1-year and 6-month maturities only are shown. Note that this filtering does not remove the annual effect that has been described earlier, with the F-statistics confirming that there are statistically significant differences in the deviations from CIP equilibrium over the 25-year sample period.

(Insert Table 5)

The estimates of the local Hurst statistic for n= 22 and 66 and for the 6-month and 1-year CIP maturities are provided in Table 5. This Table shows the results for both the filtered and unfiltered series. Filtering removes short order autocorrelation, which is associated with short-term dependence (Lo, 1992). The Hurst coefficient provides a measure of the degree of dependence in the market and may also proxy as a measure of efficiency or departure from a random walk, which has a Hurst coefficient of H=0.5. The first finding is that there are statistically significant differences in the Hurst exponent over the sample period as revealed by the F-statistic of differences in yearly means.

The second finding is that filtering for AR(1) appears to make the series more negatively dependent. For example, the 6-month unfiltered average Hurst for the entire sample period is 0.50521 compared with 0.4708 in the filtered series. A negatively dependent series has a greater probability of a +/- change in price being following by the opposite (-/+) in the subsequent period. This would be consistent with the actions of arbitrageurs whose attempts to profit from a CIP deviation cause the disequilibrium to reverse.

(insert Figures 4, 5, 6)

The final figures highlight the contribution of short-term dependence (as shown by the AR(1) filtered series) to the overall finding of negative long-term dependence. Figure 4 plots the average annual local Hurst (22) exponent estimated on 1-year CIP residuals, while Figure 5 plots the average annual local Hurst (66) exponent estimated on 6-month CIP residuals for years 1983 to 2005. Note that the short-term contribution is plotted in Figure 6 as the difference between the unfiltered and filtered annual averages. Importantly this last figure highlights the time varying nature of both long and short-term dependence in the various series, which appears to coincide with periods of market turbulence (Taylor, 1987) such as the Asian Crisis and other key economic events such as the Iraqi Wars. This last point offers the possibility of further empirical investigation.

5. Concluding Remarks

The assumption that equilibrium prices exist between spot and forward exchange rates and interest rate markets is an important relationship in international finance. More specifically, covered interest parity (CIP) arbitrage ensures that actual prices in forward currency markets are in equilibrium with those implied by those prices estimated from the underlying interest rate differentials.

This study investigates the sensitivity over time of the difference between estimated and actual prices in forward markets. Using a daily time series from 1983 to 2005 of spot and forward USD/Yen prices for 1, 3, 6-month and 1-year maturities and the equivalent maturity short term US and Japanese interest rates, we find evidence of considerable variation in CIP deviations from equilibrium over the sample period, which tend to be one way and favour those with the ability to borrow US dollars. This result is consistent with some earlier studies, specifically Poitras (1988) and cannot be simply explained as the consequence of a transaction band. When investigated from a monthly perspective there is also statistically significant evidence of seasonality - for example, there are more arbitrage possibilities available during the northern summer than in the northern winter, which is clearly associated with the liquidity of foreign exchange markets. Importantly this study is the first to identify that there has been a significant decline in the deviations from CIP in recent years – they appear to have been almost eliminated by 2000. We attribute this to the effects of electronic trading such as the Reuters D2000 and EBS trading platforms and connected product-pricing systems, which have improved the efficiency of foreign exchange markets.

To provide an insight into the underlying causes for the deviation from CIP, regression analysis using an autoregressive and GARCH specification is undertaken on the changes in the CIP residuals. This analysis reveals that interday negative changes in spot exchange rates, positive changes in US interest rates and negative changes in yen interest rates generally matter more than changes in interday volatility in these assets. These results are maintained during different subperiods of the 25-year sample period.

Furthermore, in order to examine the long-term dependence of CIP deviations over time we employ techniques from fractal geometry after accommodating underlying autoregressive behaviour. These tests, using a local Hurst exponent, reveal episodes of both positive and negative dependence over the various sample periods, which appear to be linked to episodes of dollar decline/yen appreciation, or vice versa. Importantly, the series are generally negatively dependent, which is consistent with the actions of arbitrageurs successfully maintaining the long-term CIP equilibrium. Nonetheless attention should be drawn to those periods –such as during the Asian Crisis – when markets failed to immediately reach equilibrium.

We conclude that while CIP arbitrage opportunities persisted in the yen forward market for many years – the likely effect of transaction costs and market segmentation – these opportunities have diminished and notably since 2000 almost disappeared. The recent use of electronic trading platforms and pricing in real time of equilibrium prices appears to have removed the scale and scope of earlier CIP arbitrage possibilities. Nonetheless given the time varying nature of these deviations from equilibrium the sample period under investigation remains a critical issue when investigating the presence of long-term dependence.

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Table 1. Annual Variation from CIP equilibrium for specific maturities on the USD/Yen for each year from 1983 to 2005

Year	No. of days	μ- 1 month	σ- 1 month	μ -3 month	σ 3 month	μ -6 month	σ 6 month	μ -1 year	σ 1year
1983	59	-0.0260	0.0377	-0.0443	0.0613	0.0020	0.5332	-0.1685	0.2351
1984	260	-0.0571	0.1676	-0.0818	0.1447	-0.0955	0.1652	-0.2086	0.2786
1985	261	-0.0428	0.2074	-0.0638	0.2076	-0.0794	0.2356	-0.1349	0.3266
1986	261	-0.0048	0.2162	-0.0275	0.1821	-0.0154	0.2140	-0.0259	0.2560
1987	261	-0.0447	0.1585	-0.0601	0.153	-0.0618	0.1676	-0.0688	0.2921
1988	260	-0.0166	0.1706	-0.0267	0.1570	-0.0415	0.1617	-0.0793	0.2531
1989	260	-0.0210	0.1765	-0.0316	0.1709	-0.0342	0.1428	-0.0548	0.1764
1990	261	-0.0019	0.1317	-0.0099	0.1433	-0.0171	0.1101	-0.0373	0.1470
1991	261	-0.0100	0.1537	-0.0173	0.1505	-0.0214	0.1531	-0.0702	0.1676
1992	262	-0.0108	0.1183	-0.0172	0.096	-0.0335	0.1123	-0.0757	0.1348
1993	261	-0.0118	0.1473	-0.0198	0.0888	-0.0345	0.0834	-0.0625	0.1041
1994	260	0.0074	0.1441	-0.0038	0.1240	-0.0351	0.0612	-0.074	0.0966
1995	260	-0.0077	0.1105	-0.0348	0.0945	-0.0662	0.0745	-0.1243	0.1265
1996	261	-0.0425	0.1314	-0.0491	0.044	-0.0851	0.0639	-0.155	0.1087
1997	261	-0.0747	0.5496	-0.0432	0.0364	-0.0748	0.0940	-0.1457	0.3802
1998	261	-0.0788	0.6300	-0.0759	0.1366	-0.1414	0.1961	-0.2172	0.3138
1999	261	-0.0054	0.1652	-0.028	0.1717	-0.0694	0.1928	-0.1360	0.1902
2000	259	-0.0332	0.1768	-0.0497	0.2014	-0.0746	0.2141	-0.1347	0.2107
2001	261	-0.0276	0.1785	-0.0347	0.1868	-0.055	0.2058	-0.1031	0.2096
2002	261	-0.0089	0.0658	-0.0199	0.0646	-0.0393	0.0701	-0.0817	0.0819
2003	261	-0.0008	0.0416	-0.0197	0.2348	-0.0157	0.0441	-0.0504	0.0631
2004	262	-0.0006	0.0068	-0.0039	0.0102	-0.0154	0.0175	-0.0587	0.0658
2005	69	0.0020	0.0137	-0.0038	0.0122	-0.0246	0.0266	-0.0696	0.0405
F-test (p-value)		2.81 (0.00)		5.79 (0.00)		10.84 (0.00)		15.47 (0.00)	
Sample μ		-0.0234		-0.0339		-0.0517		-0.1003	
Sample σ		0.2298		0.1454		0.1578		0.2152	
Sample skewness		0.86		-2.18		3.89		2.27	
Sample kurtosis		114.61		105.12		100.60		85.75	

Table 2. A Calendar Effect: deviations from CIP for specific maturities on the USD/Yen for each month from 1983 to 2005

Year	Number of days	μ- 1 month	σ- 1 month	μ -3 month	σ 3 month	μ -6 month	σ 6 month	μ – 1 year	σ 1 year
January	488	-0.0146	0.2186	-0.015	0.1875	-0.024	0.1739	-0.0915	0.2242
February	440	0.0109	0.1915	-0.0107	0.1526	-0.033	0.1542	-0.0871	0.1867
March	486	-0.0224	0.1561	-0.0281	0.1081	-0.0554	0.1358	-0.0885	0.1941
April	456	-0.0356	0.1813	-0.0489	0.2365	-0.0743	0.1547	-0.1129	0.1942
May	466	-0.0384	0.2101	-0.0379	0.1294	-0.0651	0.1261	-0.0962	0.1967
June	448	-0.0096	0.3962	-0.0437	0.1349	-0.0572	0.1434	-0.0771	0.3138
July	467	-0.0551	0.2825	-0.0477	0.1686	-0.0815	0.1916	-0.1168	0.2119
August	464	-0.0072	0.2705	-0.0382	0.0973	-0.06	0.11	-0.097	0.1844
September	449	-0.0119	0.1208	-0.0295	0.1117	-0.0367	0.1257	-0.0973	0.1823
October	482	-0.0339	0.2696	-0.0474	0.1412	-0.0463	0.2387	-0.1105	0.2143
November	471	-0.0138	0.1505	-0.0394	0.1105	-0.0468	0.1512	-0.1128	0.2512
December	487	-0.0439	0.1697	-0.0212	0.0984	-0.0415	0.1322	-0.1137	0.193
F-test		3.05		3.74		5.48		1.67	
p-value		0.00		0.00		0.00		0.00	

Table 3: Regression analysis of interday changes in CIP residuals across various time periods 1983 to 2005

		Explanatory variables							Autoregressive Terms			GARCH (1,1)		
Whole	c	Δe_s	$\Delta i^{\$}$	$\Delta i^{¥}$	$\Delta\sigma(e_s)$	$\Delta\sigma(i^{\$})$	$\Delta\sigma(i^{¥})$	$\Delta\delta_{t,t-1}$	$\Delta\delta_{t-1,t-2}$	$\Delta\delta_{t-2,t-3}$	a	ε^2_{t-1}	σ^2_{t-1}	AR^2
1month	-0.004 (0.000)	-0.041 (0.000)	0.127 (0.000)	-	-0.029 (0.003)	-0.036 (0.000)	-0.001 (0.064)	-0.461 (0.000)	-0.093 (0.000)	-0.240 (0.000)	0.005 (0.000)	1.870 (0.000)	0.056# (0.000)	0.23
3month	-	-0.029 (0.000)	0.169 (0.000)	-0.204 (0.000)	-	-	-0.001 (0.000)	-0.554 (0.000)	-0.266 (0.000)	-0.140 (0.000)	0.005 (0.000)-	0.558 (0.000)	0.512 (0.000)	0.30
6month	0.004 (0.001)	-0.061 (0.000)	0.338 (0.000)	-0.487 (0.000)	-	-	-0.001 (0.081)	-0.372 (0.000)	-0.171 (0.000)	-0.093 (0.000)	0.009 (0.000)-	0.304 (0.000)	0.658 (0.000)	0.32
1year	-	-0.017 (0.000)	0.696 (0.000)	-0.941 (0.000)	-0.095 (0.065)	-	-	-0.450 (0.000)	-0.243 (0.000)	-0.105 (0.000)	0.011 (0.000)-	0.183 (0.000)	0.926# (0.000)	0.40
Sample 1														
1month	-	-0.042 (0.000)	-	-0.046 (0.092)	-0.123 (0.004)	-0.014 (0.076)	-	-0.506 (0.000)	-0.326 (0.000)	-0.265 (0.000)	0.002 (0.000)	1.214 (0.000)	0.772# (0.000)	0.23
3month	-	-0.022 (0.000)	-	-0.266 (0.000)	-	-0.056 (0.017)	-	-0.429 (0.000)	-0.203 (0.000)	-0.059 (0.011)	0.013 (0.000)-	0.458 (0.000)	0.654# (0.023)	0.27
6month	-	-0.039 (0.000)	0.220 (0.001)	-0.483 (0.000)	-	-	-	-0.388 (0.000)	-0.252 (0.000)	-0.156 (0.000)	0.006 (0.000)-	0.493 (0.000)	0.922# (0.000)	0.27
1year	-	-	0.643 (0.001)	-1.257 (0.000)	-	-	-	-0.369 (0.000)	-0.221 (0.000)	-0.101 (0.000)	-	0.299 (0.000)	1.176# (0.000)	0.38
Sample 2														
1month	-	-0.040 (0.000)	0.098 (0.001)	0.056 (0.007)	-	-0.010 (0.096)	-	-0.551 (0.000)	-0.297 (0.000)	-0.147 (0.000)	0.003 (0.000)-	0.697 (0.000)	0.974# (0.023)	0.31
3month	-	-0.019 (0.000)	0.188 (0.000)	-0.053 (0.010)	-	-	-	-0.505 (0.000)	-0.255 (0.000)	-0.136 (0.000)	0.017 (0.000)-	0.600 (0.000)	0.347 (0.000)	0.29
6month	-	-0.0284 (0.000)	0.324 (0.001)	-0.377 (0.000)	-	-	-	-0.466 (0.000)	-0.258 (0.000)	-0.119 (0.000)	0.005 (0.000)-	0.844 (0.000)	0.067# (0.089)	0.26
1year	-	-0.016 (0.000)	0.667 (0.000)	-0.567 (0.000)	-	0.077 (0.002)	-	-0.440 (0.000)	-0.179 (0.000)	-0.056 (0.000)	0.023 (0.000)	0.179 (0.000)	0.628# (0.000)	0.32

Sample 3	c	Δe_s	$\Delta i^{\$}$	$\Delta i^{¥}$	$\Delta\sigma(e_s)$	$\Delta\sigma(i^{\$})$	$\Delta\sigma(i^{¥})$	$\Delta\delta_{t,t-1}$	$\Delta\delta_{t-1,t-2}$	$\Delta\delta_{t-2,t-3}$	*a*	ε^2_{t-1}	σ^2_{t-1}	AR^2
1month	-0.005 (0.000)	-0.006 (0.000)	0.005 (0.001)	-	0.023 (0.026)	0.032 (0.000)	-	-0.811 (0.000)	-0.097 (0.000)	-0.040 (0.000)	-	0.623 (0.000)	0.374 (0.000)	0.15
3month	-0.001 (0.063)	-0.009 (0.000)	0.198 (0.000)	-0.336 (0.000)	-	0.010 (0.043)	0.001 (0.000)	-0.624 (0.000)	-0.240 (0.000)	-0.174 (0.000)	-0.005	0.235 (0.000)	0.395# (0.000)	0.33
6month	0.002 (0.000)	-0.007 (0.000)	0.136 (0.000)	-0.459 (0.000)	-0.016 (0.005)	0.002 (0.055)	-0.001 (0.000)	-0.666 (0.000)	-0.461 (0.000)	-0.462 (0.000)	-	0.418 (0.000)	0.029# (0.000)	0.33
1year	-0.004 (0.000)	-0.009 (0.000)	0.417 (0.000)	-1.160 (0.000)	-0.093 (0.003)	-	-0.001 (0.000)	-0.741 (0.000)	-0.497 (0.000)	-0.254 (0.000)	-0.003	0.111 (0.000)	0.012# (0.000)	0.51
Sample 4														
1month	-	-	-	-0.120 (0.000)	-	-	-0.001 (0.031)	-0.536 (0.000)	-0.418 (0.000)	-0.300 (0.000)	-	0.220 (0.000)	0.757 (0.000)	0.33
3month	-	-0.011 (0.000)	0.138 (0.000)	-	-	0.026 (0.000)	-0.002 (0.000)	-0.482 (0.000)	-0.145 (0.000)	-0.148 (0.000)	-	0.483 (0.000)	0.362# (0.000)	0.31
6month	-	0.004 (0.007)	0.004 (0.000)	-0.620	-0.038 (0.000)	-0.003 (0.000)	-0.001 (0.000)	-0.877 (0.000)	-0.444 (0.000)	-0.257 (0.000)	-	-	1.071 (0.000)	0.29
1year	-	-0.008 (0.000)	0.442 (0.000)	-1.116 (0.000)	-0.067 (0.003)	-0.013 (0.000)	0.001 (0.000)	-0.707 (0.000)	-0.334 (0.000)	-0.172 (0.000)	-	0.032 (0.000)	0.048# (0.023)	0.33

Notes: #GARCH (2,2) needed to stabilize conditional volatility structure.

Table 4. Annual Variation in CIP Residuals after application of AR(1) filter: 1983 to 2005

Year	Number of days	μ- 1 month (AR1)	σ- 1 month (AR1)	μ- 1 month (None)	σ- 1 month (None)
1983	59	-0.0219	0.0374	-0.0260	0.0377
1984	260	-0.0482	0.1637	-0.0571	0.1676
1985	261	-0.0357	0.2042	-0.0428	0.2074
1986	261	-0.0041	0.2179	-0.0048	0.2162
1987	261	-0.0375	0.1599	-0.0447	0.1585
1988	260	-0.0139	0.1722	-0.0166	0.1706
1989	260	-0.0177	0.1753	-0.0210	0.1765
1990	261	-0.0016	0.1323	-0.0019	0.1317
1991	261	-0.0084	0.1505	-0.0100	0.1537
1992	262	-0.0091	0.1230	-0.0108	0.1183
1993	261	-0.0100	0.1490	-0.0118	0.1473
1994	260	0.0063	0.1449	0.0074	0.1441
1995	260	-0.0065	0.1120	-0.0077	0.1105
1996	261	-0.0357	0.1310	-0.0425	0.1314
1997	261	-0.0627	0.5366	-0.0747	0.5496
1998	261	-0.0657	0.6245	-0.0788	0.6300
1999	261	-0.0051	0.1572	-0.0054	0.1652
2000	259	-0.0278	0.1776	-0.0332	0.1768
2001	261	-0.0232	0.1821	-0.0276	0.1785
2002	261	-0.0075	0.0647	-0.0089	0.0658
2003	261	-0.0006	0.0421	-0.0008	0.0416
2004	262	-0.0005	0.0065	-0.0006	0.0068
2005	69	0.0017	0.0124	0.002	0.0137
F-test		2.01		2.81	
p-value		0.00		0.00	
Sample σ			0.226		0.2298

Table 5. Local Hurst Statistic estimated from the 1-year and 6-month CIP residuals for n = 22 and 66, unfiltered and AR(1) filtered

		Unfiltered CIP Residuals				AR(1) Filtered CIP Residuals			
		1-year		6-month		1-year		6-month	
Year	N	μ H22	σ H22	μ H66	σ H66	μ H22	σ H22	μ H66	σ H66
1984	260	0.4999	0.1292	0.47682	0.0525	0.4515	0.128	0.4046	0.1521
1985	261	0.5311	0.1092	0.4654	0.04598	0.4804	0.1086	0.433	0.0449
1986	261	0.5686	0.1809	0.51584	0.08036	0.527	0.1875	0.4903	0.0864
1987	261	0.5525	0.1479	0.46234	0.07323	0.498	0.1584	0.4265	0.0746
1988	260	0.5484	0.1206	0.483	0.08574	0.5038	0.1137	0.4511	0.0806
1989	260	0.5629	0.1466	0.52112	0.08609	0.5282	0.1532	0.4884	0.0875
1990	261	0.5343	0.1292	0.49795	0.07505	0.4894	0.1251	0.4742	0.0705
1991	261	0.5587	0.1335	0.55033	0.11678	0.5174	0.1353	0.5178	0.1205
1992	262	0.5838	0.1485	0.48993	0.07559	0.5354	0.1642	0.4621	0.0731
1993	261	0.575	0.1796	0.60364	0.20077	0.5202	0.1871	0.586	0.2039
1994	260	0.5398	0.094	0.53466	0.10751	0.4882	0.1	0.5038	0.1079
1995	260	0.4789	0.1161	0.54889	0.13343	0.4424	0.1213	0.5272	0.1421
1996	261	0.4986	0.0701	0.49563	0.04101	0.4472	0.0748	0.4668	0.044
1997	261	0.5528	0.206	0.41435	0.04333	0.507	0.2244	0.3849	0.041
1998	261	0.4582	0.0818	0.4307	0.0644	0.4128	0.0786	0.3893	0.0588
1999	261	0.5447	0.1289	0.52996	0.09695	0.4875	0.132	0.4915	0.0784
2000	259	0.4848	0.1434	0.48394	0.05262	0.4494	0.1459	0.4614	0.0499
2001	261	0.5293	0.1258	0.45146	0.06782	0.4827	0.1201	0.4308	0.0654
2002	261	0.4917	0.1774	0.57053	0.18171	0.4415	0.1811	0.5389	0.1832
2003	261	0.5183	0.138	0.61602	0.15653	0.465	0.1396	0.5893	0.1547
2004	262	0.4959	0.1478	0.41563	0.04347	0.447	0.1505	0.384	0.0426
2005	69	0.5182	0.0805	0.45703	0.02012	0.4573	0.0794	0.4126	0.0188
F-statistic			15.09		77.92		22.1		82.7
Local Hurst μ		0.52873		0.50251		0.4796		0.4708	
Local Hurst σ		0.139		0.09988		0.1439		0.1039	

Figure 1: The mechanics of Covered Interest Parity (CIP) arbitrage using the US$-Yen spot, forward and Euro-interest rates

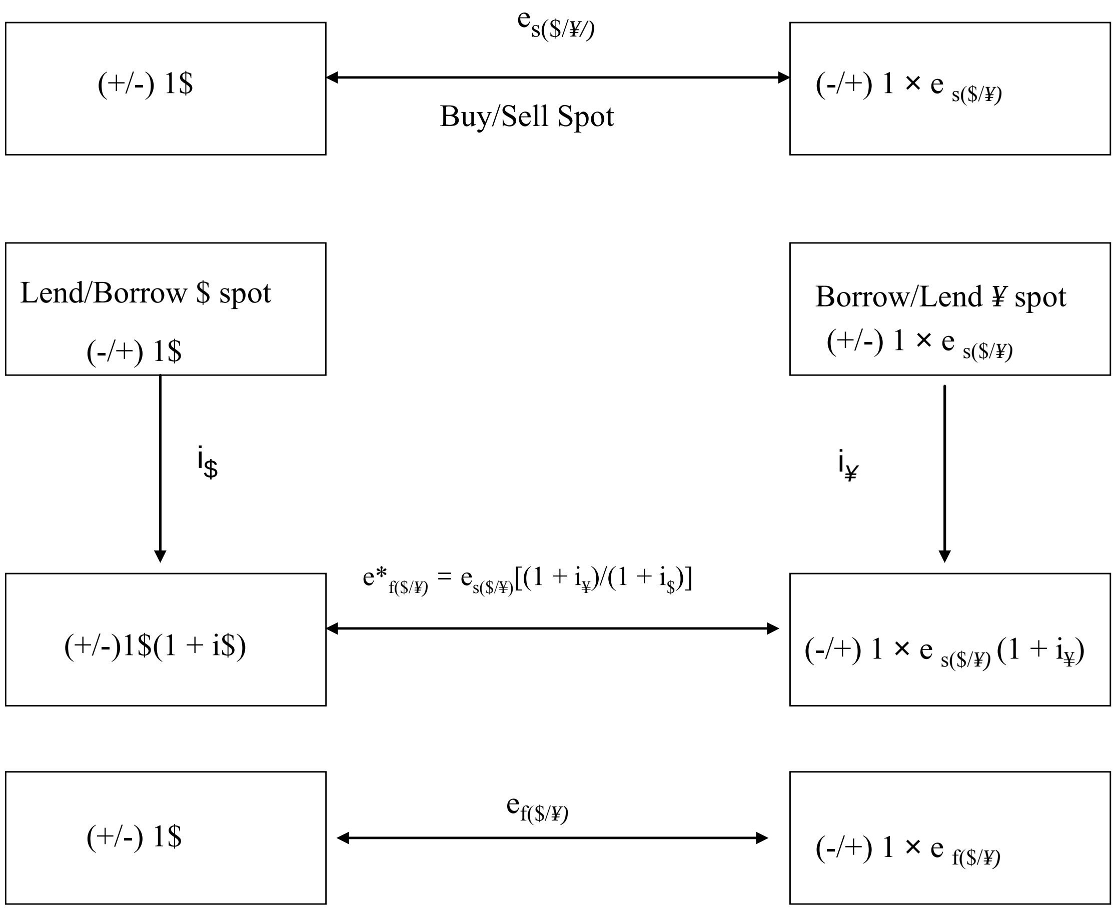


Figure 2: Plot of the monthly mean deviation from CIP equilibrium over the 25-year sample period 1983-2005 for 1, 3, 6-month and 1-year maturities. The horizontal axis is in years, while the y-axis is the deviation from CIP equilibrium in basis points.

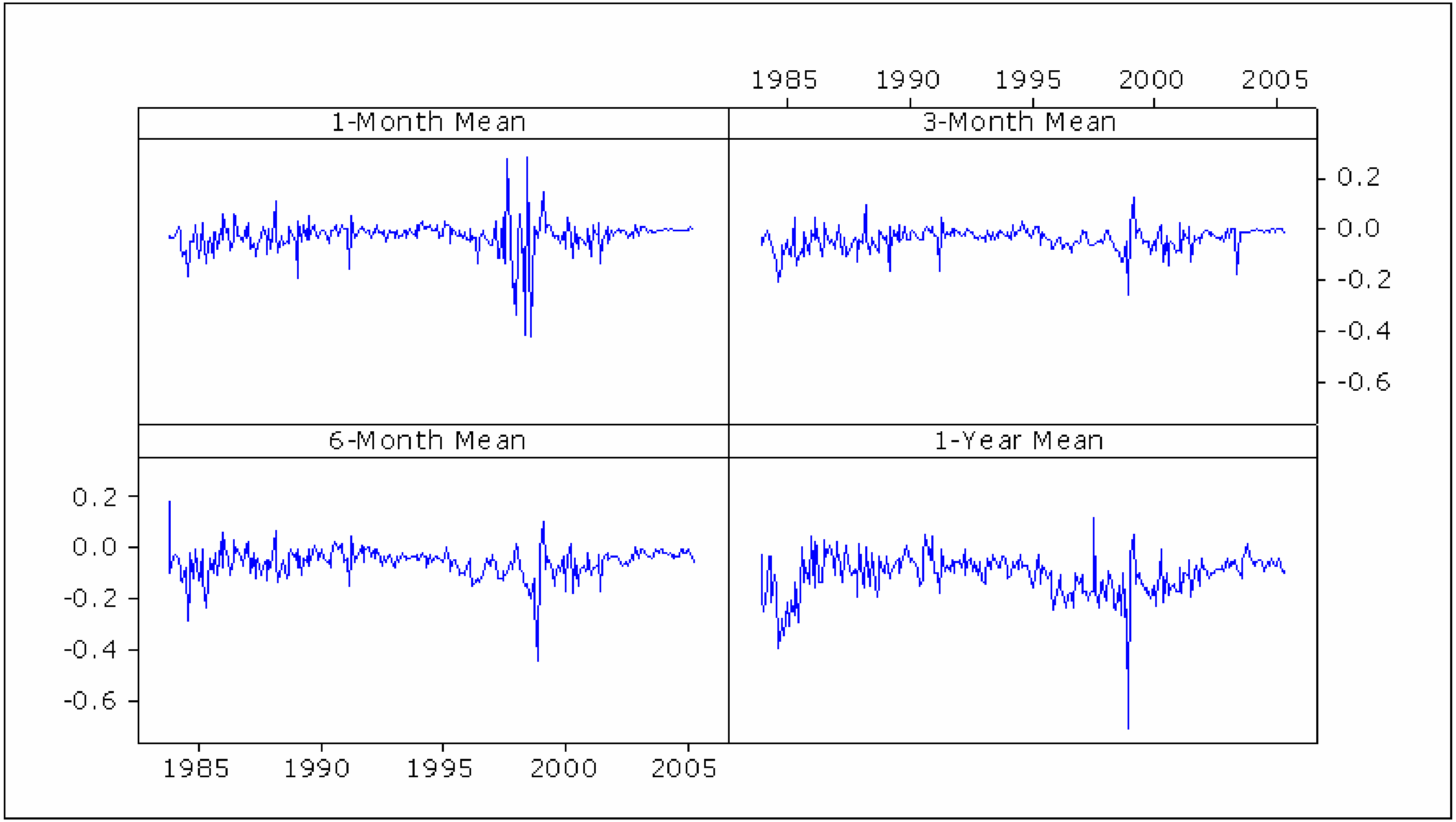


Figure 3: Plot of the SD of the monthly deviation from CIP equilibrium over the 25-year sample period 1983-2005 for 1, 3, 6-month and 1-year maturities. The horizontal axis is in years, while the y-axis is the deviation from CIP equilibrium in basis points

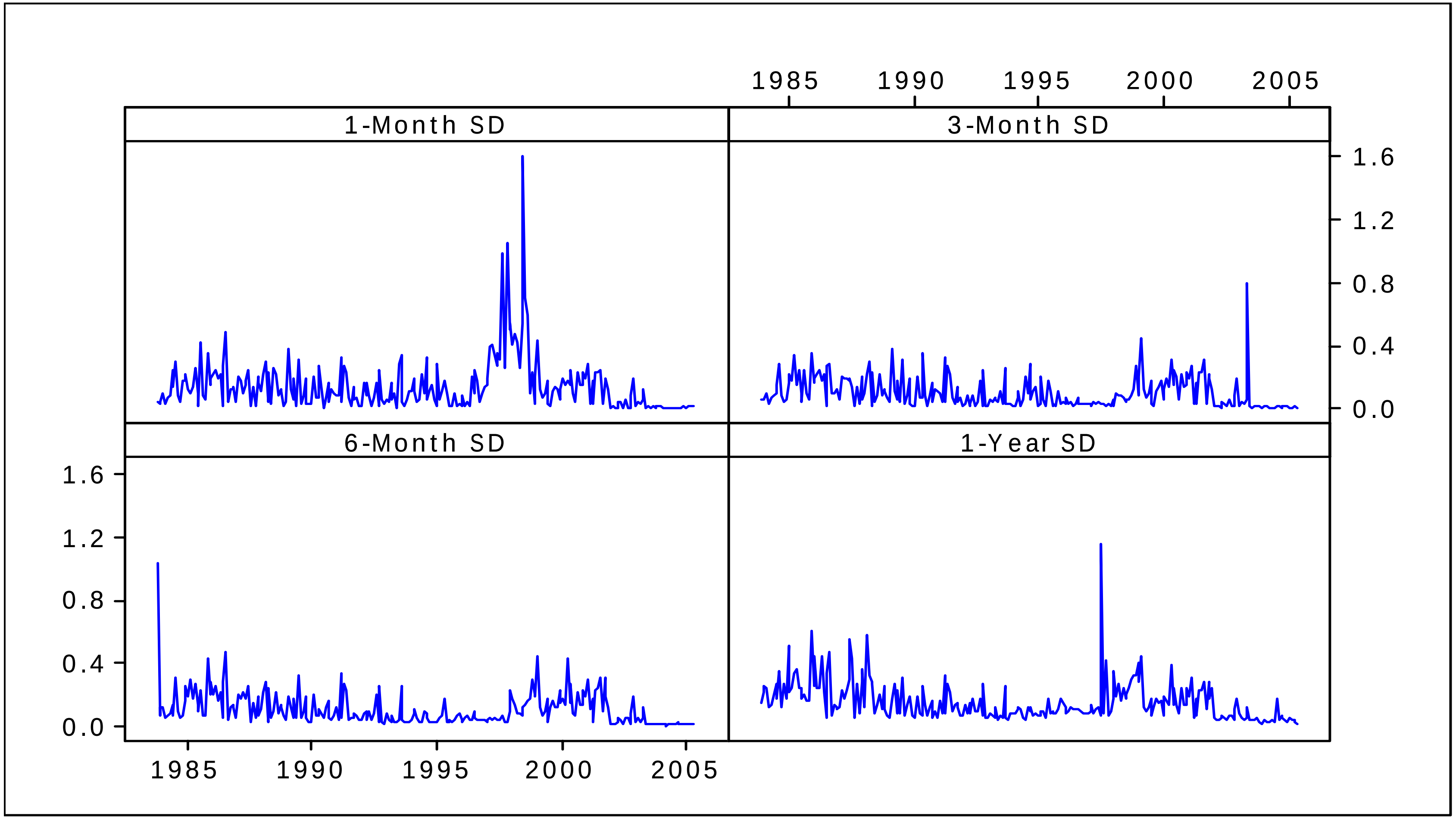


Figure 4: **Plot of average annual local Hurst (22) exponent estimated on 1-year CIP residuals for years 1983 to 2005**

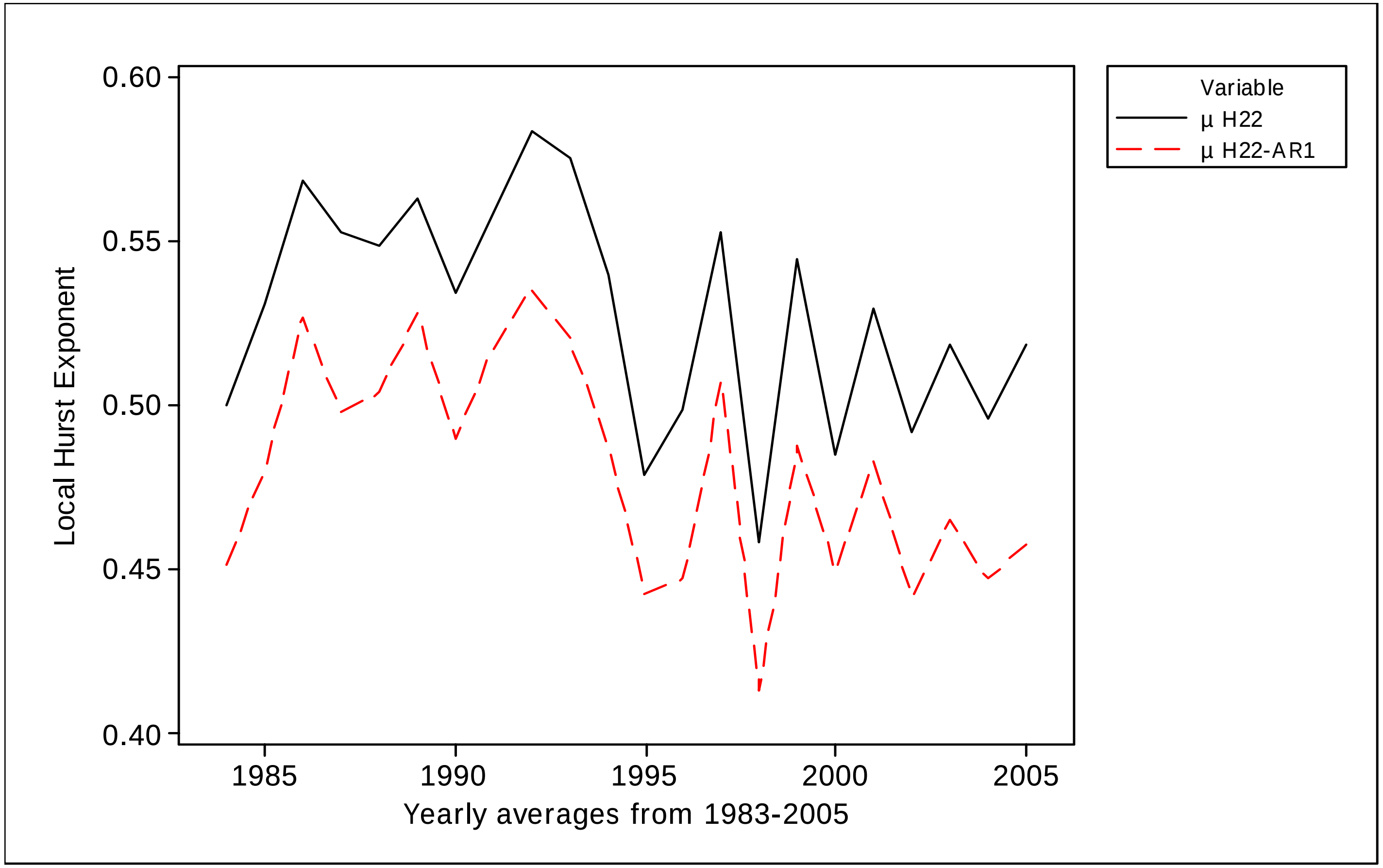


Figure 5: **Plot of average annual local Hurst (66) exponent estimated on 6-month CIP residuals for years 1983 to 2005**

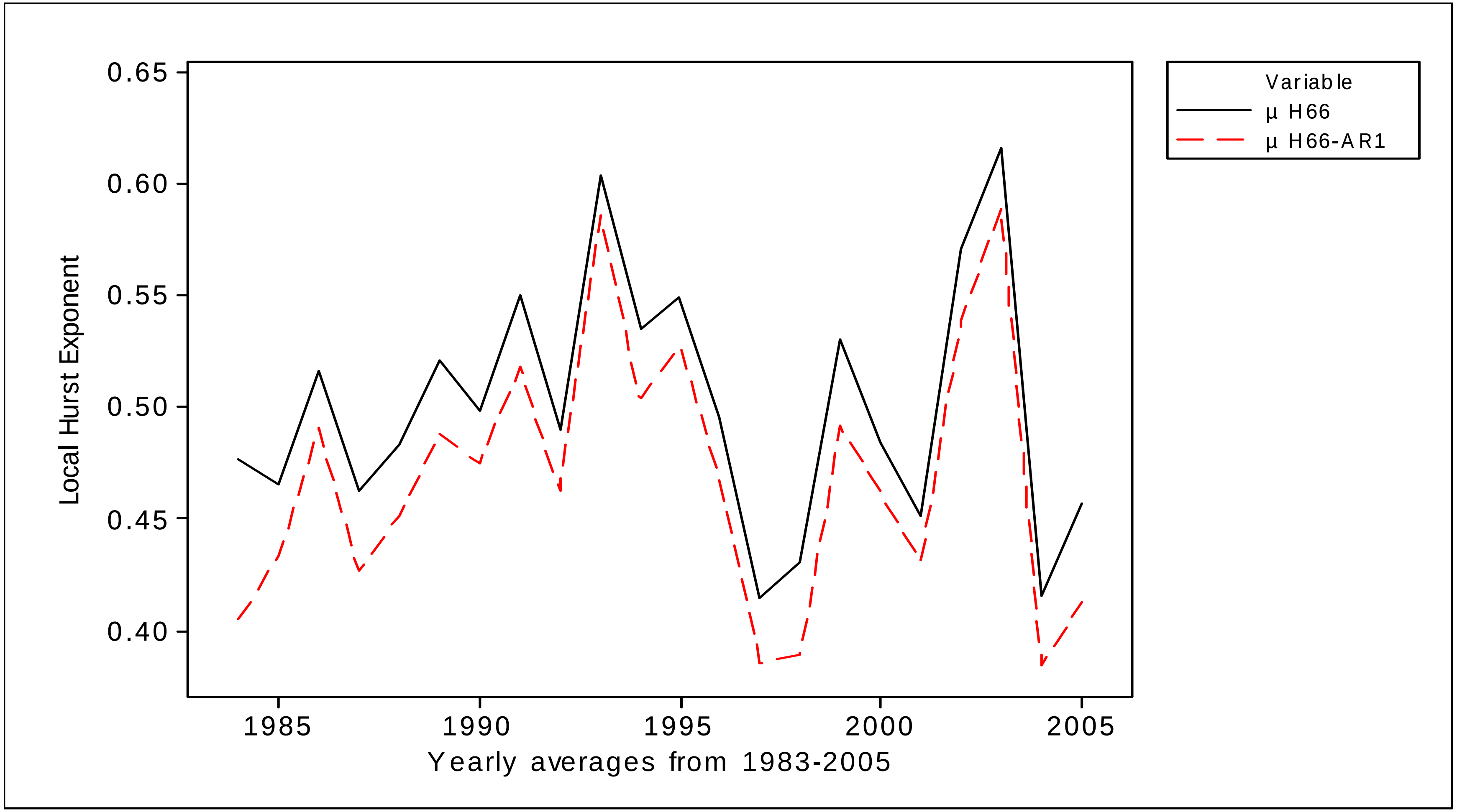
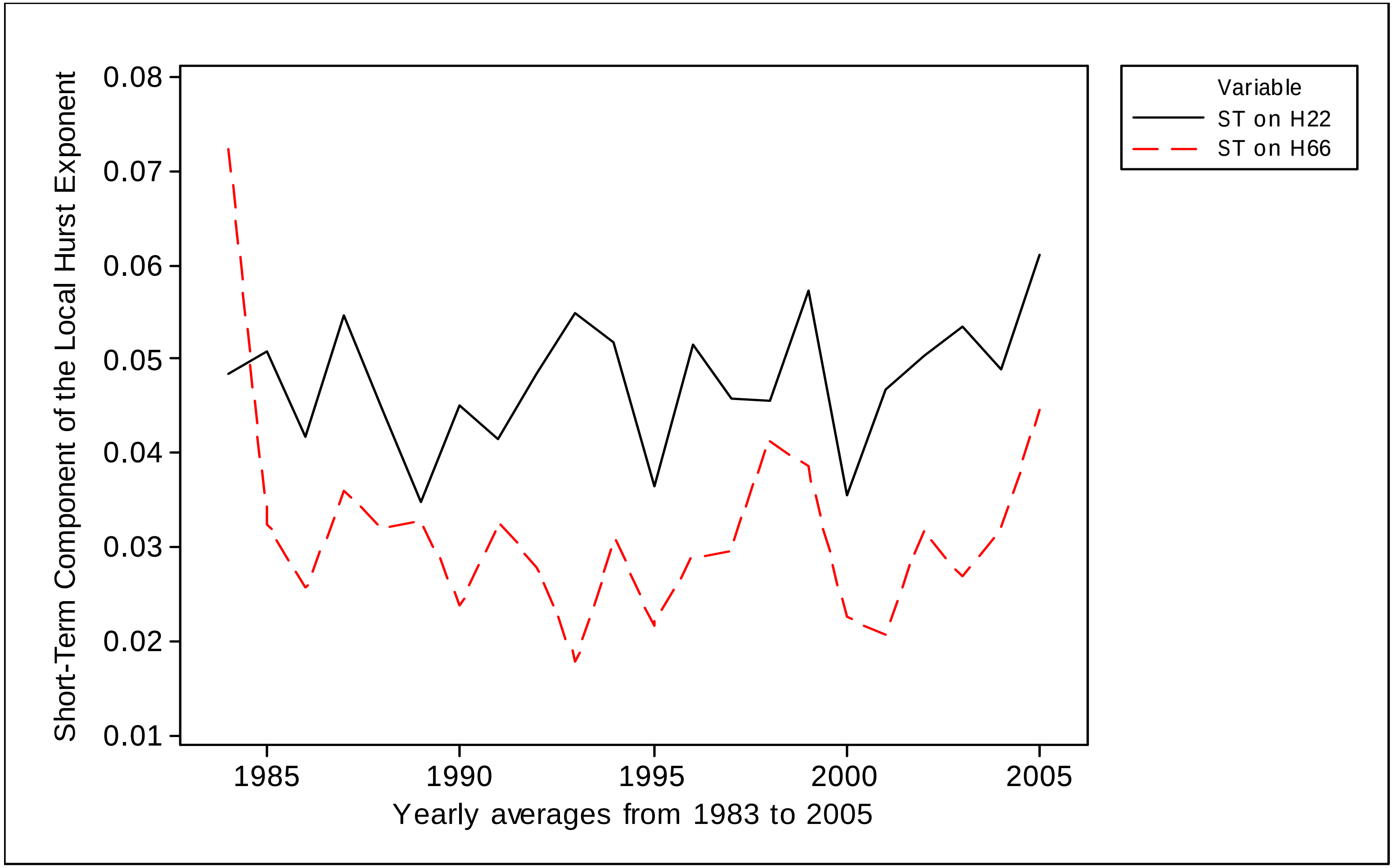


Figure 6: Contribution of Short-Term Dependence to the Local Hurst Exponent for H22 (1-year) and H66 (6 month) maturities



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