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(Citation)

神戸大学経済学研究科 Discussion Paper, 801

(Issue Date)

2008

(Resource Type)

technical report

(Version)

Version of Record

(URL)

<https://hdl.handle.net/20.500.14094/81000053>



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April 2008  
Discussion Paper No.0801

**GRADUATE SCHOOL OF ECONOMICS  
KOBE UNIVERSITY**

**ROKKO, KOBE, JAPAN**

# Intraday Evidence of the Informational Efficiency of the Yen/Dollar Exchange Rate<sup>☆</sup>

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## Abstract

The informational efficiency of the yen/dollar exchange rate is investigated in five market segments within each business day from 1987 to 2007. Among the results, we first find that the daily exchange rate has a cointegrating relationship with the cumulative price change of the segment for which the London and New York markets are both open, but not with that of any other segments. Second, the cumulative price change of the London/N.Y. segment is the most persistent among the five market segments in the medium- and long-run. These results suggest that the greatest concentration of informed traders is in the London/N.Y. segment where intraday transactions are the highest. This is consistent with the theoretical prediction by Admati and Pfleiderer (1988) that prices are more informative when trading volume is heavier.

*JEL classification:* F31

*Keywords:* Informational efficiency; Market segments; Yen/dollar exchange rate

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<sup>☆</sup> We thank the participants of the Institute of Statistical Research Conference in Barcelona for useful comments. This work is financially supported by Grant-in-Aid for Young Scientists (B18730211, K. Iwatsubo; B19730216, Y. Kitamura) and Yamada Academic Research Promotion Fund.

## 1. Introduction

Foreign exchange markets do not sleep. Traders can exchange currencies 24 hours a day in any market around the world if trading counterparties are available. Nevertheless, in reality reasonably well-defined opening and closing times do exist and the activity of each market has a typical intraday seasonality. Several recent studies exploit the Electronic Broking Services (EBS)<sup>1</sup> data to document intraday volume patterns in the foreign exchange market (Chaboud et al. (2004, 2007), Ito and Hashimoto (2006), and Cai et al. (2007))<sup>2</sup>. Ito and Hashimoto (2006) find that there is a U-shape pattern of volume – namely, heavy trading in the beginning and at the end of the trading day and relatively light trading in the middle of the day – in the Tokyo and London markets for the yen/dollar and euro/dollar rates, but not in the N.Y. market of those currencies. A more notable confirmed feature is that the overlapping business hours encourage inter-regional transactions and an overall surge in activities. For the yen/dollar transactions, for example, trading volume is the highest when the London and New York business hours overlap and the second highest when the Tokyo and London markets are both open (Figure 1). Taken together, this evidence suggests that the U-shape pattern of intraday volume results from the concentration of transactions during overlapping business hours.

\*\*\*\*\* Figure 1 around here \*\*\*\*\*

The objective of this paper is to investigate the empirical link between intraday trading volume and informational efficiency. There has been a great deal of empirical work on the volume-volatility relationship (e.g. Tauchen and Pitts (1983)), but there are few studies on the relationship between volume and informational efficiency. Our results help clarify why trading volume is high during the overlapping business hours of foreign exchange markets.

There are two polar views regarding the relationship between volume and informational efficiency: the asymmetric information view and the inventory control view.

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<sup>1</sup> In the last few years, the brokers markets have been revolutionized by the introduction of two electronic broking systems (Reuters D2002-2 and Electronic Broking Services [EBS]) which have effectively taken over from voice brokers.

<sup>2</sup> Chaboud et al. (2004, 2007) and Cai et al. (2007) find sharp spikes in volume at the times of scheduled macroeconomic data releases, spot foreign exchange fixings, and the standard expiration times of foreign exchange options.

The asymmetric information view argues that trades are more informative when trading volume is high, while the inventory control view holds that trades are less informative when trading volume is high. Theory admits both possibilities, depending on the posited information structure.

To understand the asymmetric information view, consider the model of Admati and Pfleiderer (1988). In order to minimize their losses to informed traders, discretionary liquidity traders prefer to trade when the market is thick – that is, when their trading has little effect on prices. With more liquidity trading in a given period, more informed traders tend to trade with liquidity traders. This makes it even more attractive for liquidity traders to trade in that period because competition among the informed traders reduces their total profit, which benefits the liquidity traders. Admati and Pfleiderer (1988) argue that these strong incentives for the concentration of trading create the U-shape pattern of intraday volume. An increase in the number of informed traders contributes to the informativeness of prices because they cause prices to adjust faster to information. In this situation, trading volume and the informational efficiency of prices are positively related.

On the other hand, the representative model of the inventory control view is developed by Lyons (1997). On the inventory control side, the key mechanism is hot potato trading – passing unwanted positions from dealer to dealer following an initial customer order. To clarify the argument, consider the following example. Suppose that there are ten dealers, all of whom are risk averse and currently hold zero positions. A customer sale of \$10 million worth of Japanese yen is accommodated by one of the dealers. Not wanting to carry the open position, the dealer calculates his share of this inventory imbalance – or one-tenth of \$10 million – calls another dealer, and unloads \$9 million worth of Japanese yen. The dealer receiving his trade then calculates his share of this inventory imbalance – or one tenth of \$9 million – calls another dealer, and unloads \$8.1 million worth of Japanese yen. The hot potato process continues. In the limit, the total interdealer volume guaranteed from the \$10 million customer trade is  $\$9 \text{ million} / (1-0.9) = \$90 \text{ million}$ . This simple example illustrates that the repeated passage of idiosyncratic inventory imbalances among dealers following an innovation in customer order flow can better explain the high transaction volume in foreign exchange markets.

Hot potato trading reduces the informativeness of prices. Information aggregation by dealers occurs through signal extraction applied to order flow. The greater the noise relative to the signal, the less effective signal extraction is. Passing the hot potato trades increase the noise in order flow and dilute informational content. Hence, trading volume and the informational efficiency of prices are negatively linked.

In this paper, we provide evidence that discriminates between the two views. Our paper shares the same motivations as the studies by Lyons (1995, 1996), in which real-time transaction data are used to document that trades occurring when intertransaction times are short are significantly less informative, supporting the inventory control view. In his work, intertransaction times are used as a proxy for trading volume. As Lyons (1995, 1996) admits, however, the assumption that intertransaction times are exogenous may be problematic because earlier work (e.g., Hausman, Lo and MacKinlay, 1992) has tested for and rejected the exogeneity of the length of time between transactions at a conventional significance level. The short sample period covering only five trading days of the week is another caveat (Mello, 1996).

The current paper takes a different approach to the analysis of the volume-information relationship. Taking the intraday volume pattern as exogenous, we examine the information efficiency of the intraday yen/dollar rate. Specifically, we split the foreign exchange market into five market segments (Tokyo, London-only, London/N.Y. overlap, N.Y.-only, and Pacific) and compare price informativeness among the segments.

A key feature of our analysis is the use of long-term data for a period of 20 years spanning from 1987 to 2007. This enables us to provide robust empirical evidence on informational efficiency and the volume-information relationship. Although Ito and Hashimoto (2006) reveal that intraday patterns of volume have been quite stable in recent years, in spite of the technological advances and the introduction of the electronic broking system, they may have changed gradually over the last two decades. It is a well-known stylized fact, however, that transaction volume is the highest in the segment where the London and N.Y. markets are both open (see Guillaume et al., 1995). Based on this fact, we explore the price informativeness between the market in overlapping business hours (London/N.Y.) and the rest of the market segments and consider the volume-information relationship. We believe that the endogeneity problem is partially

solved, without resorting to the use of high-frequency exchange rate and order flow data, by exploiting the stylized fact that trading volume has been higher in the London/N.Y. segment than in any other segment.

Another notable feature of this paper is the introduction of a new concept for intraday exchange rate analysis: the “cumulative price change” of market segments. The daily yen/dollar exchange rate change from  $t$  to  $t+1$  ( $\Delta P_{t+1} = P_{t+1}^{TY} - P_t^{TY}$ ), measured at the Tokyo opening, can be decomposed into the price changes of the five market segments:

$$\Delta P_{t+1} = \Delta P_t^{TY} + \Delta P_t^{LN} + \Delta P_t^{L/N} + \Delta P_t^{NY} + \Delta P_t^{PA}, \quad (1)$$

where  $\Delta P_t^{TY}$  is the price change in the Tokyo segment (TY) from its opening to closing ( $\Delta P_t^{TY} = P_t^{TY,CL} - P_t^{TY,OP}$ ),  $\Delta P_t^{LN}$  is that in the London segment (LN) from the Tokyo closing to the N.Y. opening ( $\Delta P_t^{LN} = P_t^{NY,OP} - P_t^{TY,CL}$ ),  $\Delta P_t^{L/N}$  is that in the London/N.Y. overlap (L/N) from the N.Y. opening to the London closing ( $\Delta P_t^{L/N} = P_t^{LN,CL} - P_t^{NY,OP}$ ),  $\Delta P_t^{NY}$  is that in the N.Y. segment (NY) from the London closing to the N.Y. closing ( $\Delta P_t^{NY} = P_t^{NY,CL} - P_t^{LN,CL}$ ), and  $\Delta P_t^{PA}$  is that in the Pacific segment (PA) from the N.Y. closing to the Tokyo opening ( $\Delta P_t^{PA} = P_{t+1}^{TY,OP} - P_t^{NY,CL}$ ). The superscripts *OP* and *CL* of the price  $P$  refer to the opening and closing rates, respectively. The data is described in Section 2.

Summing both sides of Equation (1) from  $t = 1$  to  $s$ , we obtain the following equation:

$$\sum_{t=1}^s \Delta P_{t+1} = \sum_{t=1}^s \Delta P_t^{TY} + \sum_{t=1}^s \Delta P_t^{LN} + \sum_{t=1}^s \Delta P_t^{L/N} + \sum_{t=1}^s \Delta P_t^{NY} + \sum_{t=1}^s \Delta P_t^{PA}. \quad (2)$$

On the left-hand side of Equation (2), the total change in the daily exchange rate is equal to  $P_{s+1} - P_1$ , that is, the daily exchange rate at  $t = s + 1$  minus its initial value  $P_1$  (constant). On the other hand, the right-hand side equals the sum of the cumulative price changes of the five market segments. The cumulative price change (CPC, hereafter) represents the aggregated price changes of each segment from  $t = 1$  to  $s$ . Equation (3) is a simple representation of Equation (2).

$$P_{s+1} - P_1 = CPC_s^{TY} + CPC_s^{LN} + CPC_s^{L/N} + CPC_s^{NY} + CPC_s^{PA}. \quad (3)$$

One of the advantages of exploiting CPCs is their nonstationarity, which allows us to measure the size of the random walk in each CPC and hence analyze (weak-form) informational efficiency. On the contrary, a caveat of using CPCs to

compare price informativeness is that CPCs are assumed to be independent from each other. If exchange rate changes in one market segment have interregional impacts, this assumption may be problematic. In the literature of exchange rate *volatility* clustering, Engle et al. (1990) find support for the meteor shower effect (interregional volatility spillover), while Baillie and Bollerslev (1990) and Melvin and Melvin (2003) report evidence for the heat wave effect (own-region volatility persistence). Cai et al. (2007) revisit this issue using the EBS data and confirm that both effects are present. However, their results based on exchange rate *returns* (not volatility) suggest that informational linkages across trading regions are so limited that their economic significance is very small. Therefore, we do not think that the assumption of independency misleads our conclusions. To check robustness, we take the interdependency of CPCs into account in the analysis of section 3-3.

Fama (1970), summarizing the idea of informational efficiency in his classic survey, writes: “A market in which prices always ‘fully reflect’ available information is called ‘efficient’.” In an informationally efficient market, price changes must be unforecastable if they fully incorporate the expectations and information of all market participants. In practice, since it is almost impossible to take traders’ information and expectations into consideration, much research on the market efficiency hypothesis tests the forecastability of prices based only on past price changes. Among various versions of the weak-form informational efficiency, many researchers investigate the random walk hypothesis. We test this hypothesis with variance ratio tests which exploit the important property of the random walk process that the variance of random walk increments must be a linear function of the time interval.

The second hypothesis is that, in an informationally efficient market, the CPCs contribute significantly to the long-run trend of exchange rates. If the CPCs are informational efficient, they must have a stable relationship with the daily exchange rate in the long run. We explore whether there is any cointegration relationship between the daily exchange rate  $P_{s+1}$  and the CPC of each segment with careful consideration of structural breaks.

Among the results, we first find that the daily exchange rate has a cointegration relationship with the CPC of the overlapping business hours in the London/N.Y. segment, but not with any other segments. Second, the CPC of the London/N.Y. segment is



the most persistent among all five segments. The evidence indicates that the price change of the London/N.Y. segment is the most informationally efficient. From a long-term perspective, this evidence supports the asymmetric information view that trades are more informative when trading volume is high.

The organization of this paper is as follows. Section 2 briefly explains the data and basic statistics. Section 3 shows several empirical results and Section 4 concludes.

## 2. Data

In order to examine the informational efficiency of the intraday yen/dollar rate, we identify five separate periods for classifying regional hours: Tokyo, London-only, London/N.Y. overlap, N.Y.-only, and Pacific. We obtain the Tokyo opening and closing rates, the London closing rate, and the N.Y. opening and closing rates from the Nikkei Needs Financial Quest system.<sup>3</sup> The sample period runs from September 14, 1987 to May 31, 2007, based on data availability. Missing data, due to national holidays, are replaced with the closest data from the previous segment. The timing of the market segments is as follows: the Tokyo opening and closing rates are recorded at 9:00 a.m. and 17:00 p.m. in local time, respectively. Since the London opening is unavailable in our long-run data set but is very close to the Tokyo closing<sup>4</sup>, we adopt the Tokyo closing as a proxy for the London opening. The London market closes at 16:00 p.m.<sup>5</sup>, and the N.Y. market opens at 8:30 a.m. and closes at 17:00 p.m. in local time.

\*\*\*\*\*Table 1 around here\*\*\*\*\*

Table 1 displays the time schedule for the five market segments. We take into account differences in the duration of each segment due to daylight saving time. Foreign exchange trading starts in the Tokyo market, and the London market opens just after the

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<sup>3</sup> The Nikkei Needs database collects the mid-points of TTS (Telegraphic Transfer Selling rate) and TTB (Telegraphic Transfer Buying rate).

<sup>4</sup> Melvin and Melvin (2003) and Cai et al. (2007) identify the period in which Tokyo and London business hours overlap, while we include the overlap in the Tokyo segment. By their definition of the London opening time, however, the length of the Tokyo/London overlap is short, 1.5h in normal time (2.5h in DST)

<sup>5</sup> Since the WM/Reuters spot foreign exchange fixing occurs at 16:00 p.m. in London, the 16:00 p.m. FX rate is commonly used. We are aware that 16:00 p.m. is a little earlier than the actual closing time of the London market. This may lead to a shorter London/N.Y. segment than the actual. Hence, our conservative definition of the segment does not give misleading results.

Tokyo market closes. Since the N.Y. market opens before the London closes, active trading takes place when both markets are open for several hours. The N.Y. market closes several hours after the London market. By our definition, the Pacific market (e.g., Sydney) is very limited, starting from the N.Y. closing to the Tokyo opening on the next day. Although this short period is included in the Asia segment in the analysis of Cai et al. (2007), as shown in Figure 1, the yen/dollar trading in this market segment is the least active in a day.

\*\*\*\*\*Table 2 around here\*\*\*\*\*

Table 2 presents the summary statistics on price changes for the five market segments. The per hour variance of price changes for each segment are shown, and is defined as the variance adjusted for the number of trading hours in each market segment. To correct for differences in segment duration, we divide the price change ( $\Delta P_t^i$ ) by the square root of business hours in segment  $i$  on date  $t$  and calculate its variance. The fifth column of Table 2 reports the variance ratio of the returns of London/N.Y. to those of each segment. Simple variance ratio tests reveal that the London/N.Y. segment is far more volatile than any other segment. This is consistent with Ito and Roley's (1987) argument that the high volatility on late London and early N.Y. business hours may reflect a great deal of news coming from the European and N.Y. markets.

### 3. Empirical Analysis

#### 3.1. Cointegration tests

Figures 2-1 through 2-5 show the CPC for each segment and the daily exchange rate throughout the sample period. We adopt the Tokyo opening rate at  $t+1$  as the daily yen/dollar rate. In Figure 2-3, we observe a long-run stable relationship between the daily exchange rate and the London/N.Y. CPC. It is also noteworthy that the Tokyo CPC shows a steadily depreciation in the dollar since 2000, while the N.Y. CPC shows a constant appreciation of the dollar during the same period. This sharp contrast is puzzling, but we limit ourselves to pointing out this phenomenon.

\*\*\*\*\*Figures 2 around here\*\*\*\*\*

We next conduct empirical tests on the stability of the relationship between the daily yen/dollar rate and the CPC. The stable relationship suggests that exchange rate pricing in the market segments greatly contributes to the determination of long-run ex-

change movements.

To test these hypotheses, we regress the daily yen/dollar rate (the Tokyo opening rate at  $t+1$ ), onto each CPC at  $t$  and perform residual-based cointegration tests. Before the cointegration analyses, we confirm the nonstationarity of each variable using ADF unit root tests. The results, presented in Table 3, indicate that all variables have a unit root in levels but not in first differentials. This indicates that they are all integrated of order one ( $I(1)$ ).<sup>6</sup>

\*\*\*\*\*Table 3 around here\*\*\*\*\*

\*\*\*\*\*Table 4 around here\*\*\*\*\*

The results of the cointegration test (ADF-type test) in Table 4 suggest that a cointegration relationship is found only between the daily yen/dollar rate and the London/N.Y. CPC. This result is consistent with the evidence that appears in Fig 2., but should be treated carefully because ADF-type tests have a low power of rejection for the unit root hypothesis in stationary variables with a structural break. An alternative test is to consider the existence of structural breaks. Gregory and Hansen (1996a, b) developed a cointegration test which allows for the possibility of a regime shift. The virtue of their test is that the structural break point is endogenously determined rather than being assumed to be predetermined. Following Gregory and Hansen (1996a), we consider the following three types of regression models:

(1) “Level shift” model (Model C)

$$P_{t+1} = \alpha + \delta_{\alpha} D_t + \beta CPC_t^i + v_{t+1}, \quad (4)$$

(2) “Regime shift” model (Model C/S)

$$P_{t+1} = \alpha + \delta_{\alpha} D_t + \beta CPC_t^i + \delta_{\beta} D_t CPC_t^i + v_{t+1}, \quad (5)$$

(3) “Regime and trend shift” model (Model C/S/T)

$$P_{t+1} = \alpha + \delta_{\alpha} D_t + \beta CPC_t^i + \delta_{\beta} D_t CPC_t^i + \gamma Trend + \delta_{\gamma} D_t Trend + v_{t+1}, \quad (6)$$

where  $D_t$  is a dummy variable with a value of unity after the break point period and zero

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<sup>6</sup> The ADF test statistic is calculated with the assumption of whiteness of residuals (Hamilton (1994)). Although the *Lag 1* ADF test statistic for the *level* of the daily rate is statistically significant at 10%, this result is dubious since the Portmanteau statistic [Q(10)] for its residuals rejects no serial correlations at the 1% significance level (Q[10]=31.015).

otherwise.  $\nu_{t+1}$  is a disturbance. Equation (4) allows for an endogenous break in the intercept and Equation (5) both in the intercept and the slope. In addition, Equation (6) allows for an endogenous break in the intercept, the slope and the deterministic trend. In searching for endogenous breakpoints, we compute the Phillips  $Z_t^*$  test statistics for each possible regime shift, take the smallest value of  $Z^*$  and compare it to the corresponding critical value of the null. When the null is rejected, one structural break in a cointegration vector is found at that point (Gregory and Hansen (1996a)).

\*\*\*\*\*Table 5 around here\*\*\*\*\*

Table 5 reports possible structural break dates and the minimum Phillips  $Z_t^*$  test statistics for Equations (4)-(6). In the London/N.Y. segment, the statistics of the “Regime and trend shift” (Model C/S/T) indicate a cointegration relationship between the CPC and the spot rate. For this segment, we divide the sample at the breakpoint and examine whether there are more breakpoints in each subsample, but no more breakpoints are found.<sup>7</sup>

In the London/N.Y. segment, both the ADF and the Gregory and Hansen tests suggest cointegration. In such a case, Gregory and Hansen (1996a) recommend not to infer that a structural change in the cointegration vector has occurred since the later test is more powerful for the detection of the existence of cointegration than the ADF. Therefore, we conclude that there is a stable relationship between the daily exchange rate and the CPC of the London/N.Y. segment with no structural breaks in the corresponding cointegration vector. The results in Fig 2 and Table 4 to 5 suggest that the London/N.Y. overlap contributes significantly to the long-run trend of the daily yen/dollar rates.

### 3.2. *Random Walk Tests*

In this sub-section, we test the random walk hypothesis for the CPC of each segment. In doing so, we employ Lo and MacKinlay’s (1988) variance-ratio analysis and Wright’s (2000) rank- and sign-based variance ratio tests to examine the size of the random walk in intraday exchange rates.

The variance ratio test exploits the fact that the variance of the increments in a

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<sup>7</sup> The results are not displayed in the paper but are available upon request.

random walk is linear in the sampling interval; that is, if a series follows a random walk process, then the variance of the  $k$ -period return would be  $k$  times the variance of the one-period return. In this study the variance-ratio at lag  $k$ ,  $VR_k^i$ , is defined as

$$VR_k^i = \frac{\frac{1}{Tk} \sum_{t=k+1}^T (y_t^i + y_{t-1}^i + \dots + y_{t-k}^i - k\hat{\mu}^i)^2}{\frac{1}{T} \sum_{t=k+1}^T (y_t^i - \hat{\mu}^i)^2} \quad i=1(TY), \dots, 5(PA), \quad (7)$$

where  $y_t (= CPC_t - CPC_{t-1})$  is the exchange rate differential at a daily interval and  $\hat{\mu}^i = T^{-1} \sum_{t=1}^T y_t^i$ . Under the random walk hypothesis, the variance ratio  $VR_k^i$  equals one for any  $k$  chosen. Lo and MacKinlay (1988) proposed the heteroskedasticity-robustified statistic:

$$M^i = (VR_k^i - 1) \left( \sum_{j=1}^{k-1} \left[ \frac{2(k-j)}{k} \right]^2 \delta_j^i \right)^{-\frac{1}{2}}, \quad (8)$$

where

$$\delta_j^i = \left\{ \sum_{t=j+1}^T (y_t^i - \hat{\mu}^i)^2 (y_{t-j}^i - \hat{\mu}^i)^2 \right\} / \left[ \sum_{t=1}^T (y_t^i - \hat{\mu}^i)^2 \right]^2.$$

Wright (2000) proposes alternative variance-ratio tests that use the ranks and signs of returns. The rank- and signed-based variance ratio tests have several advantages: (1) they may be more powerful than other tests if the returns are not normally distributed, and (2) it is possible to yield their exact distributions. The rank-based variance ratio tests are exact under the assumption that returns are iid, and the signed-based tests are exact even in the presence of conditional heteroskedasticity.

Let  $r^i(y_t^i)$  be the rank of  $y_t^i$  among  $y_1^i, y_2^i, \dots, y_T^i$  for  $i=1(TKO), \dots, 5(PAC)$ . Define  $r_t^i = \Phi^{-1}(r^i(y_t^i)/(T+1))$  where  $\Phi$  is the standard normal cumulative distribution function. The rank-based variance ratio tests substitute  $r_t^i$  in place of  $y_t^i$  in the definition of the test statistic  $M^i$  (eq. 8), noting that the rank is standardized to have a mean of zero. The test statistic is

$$R_k^i = \left( \frac{\frac{1}{Tk} \sum_{t=k+1}^T (r_t^i + r_{t-1}^i + \dots + r_{t-k}^i)^2}{\frac{1}{T} \sum_{t=1}^T (r_t^i)^2} - 1 \right) \times \left( \frac{2(2k-1)(k-1)}{3kT} \right)^{\frac{1}{2}}, \quad i=1(TY), \dots, 5(PA). \quad (9)$$

Under the assumption that  $y_t^i$  is iid,  $r^i(y_t^i)$  is a random permutation of the numbers 1, 2, 3, ..., T, each with equal probability, giving the distribution of the test statistic.

By using the signs rather than ranks, the sign-based variance ratio tests are defined. Assume that the mean of  $y_t^i$ ,  $\hat{\mu}^i$ , is 0, which is quite a reasonable assumption for intraday exchange rate differentials.  $s^i(y_t^i)$  is equal to 1 with probability 0.5 (when  $y_t^i \geq 0$ ) and is equal to -1 otherwise. Define the variance ratio statistic as

$$S_k^i = \left( \frac{\frac{1}{Tk} \sum_{t=k+1}^T (s_t^i + s_{t-1}^i + \dots + s_{t-k}^i)^2}{\frac{1}{T} \sum_{t=1}^T (s_t^i)^2} - 1 \right) \times \left( \frac{2(2k-1)(k-1)}{3kT} \right)^{-\frac{1}{2}}, \quad i = 1(TY), \dots, 5(PA). \quad (10)$$

\*\*\*\*\*Table 6 around here\*\*\*\*\*

Table 6 presents each CPC's variance ratio statistics, Lo and MacKinlay's (1988) heteroskedasticity-robust standard normal z-statistics, Wright's (2000) nonparametric rank- and signed-based test statistics for lags ( $k$ ) of 5, 10, 20, and 30. As shown in Table 6, the CPC of the Tokyo and Pacific segments are serially dependent, while those of London, London/N.Y. and N.Y. segments are serially independent. This result holds if we extend the lags to 200 ( $k=200$ ). Indeed the variance ratio tests are powerful in measuring the random walk in each CPC, but we cannot infer which is more informationally efficient among the CPCs of London, London/N.Y. and N.Y. We compare these in the next subsection.

### 3.3. Mean Squared Error Decomposition

In order to measure the contribution of movements in each CPC in determining the overall movement of the daily exchange rate, we exploit Engel's (1999) mean-squared error (MSE)<sup>8</sup> decomposition. The MSE of the change in the exchange rate – which is the sum of the squared drift and the variance – is a comprehensive measure of movement. Following Engel (1999), we calculate two different methods for measuring the fraction of the MSE of the exchange rate change accounted for by the MSE of each

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<sup>8</sup> The MSE is defined as

$MSE(CPC_t - CPC_{t-k}) = \text{var}(CPC_t - CPC_{t-k}) + [\text{mean}(CPC_t - CPC_{t-k})]^2$ , where “mean” and “var” are mean and variance, respectively.

CPC change. The first decomposition is

$$\frac{MSE(CPC_t^i - CPC_{t-k}^i)}{\sum_i MSE(CPC_t^i - CPC_{t-k}^i)} \quad i = 1(TY), \dots, 5(PA), \quad (11)$$

and the second is

$$\frac{MSE(CPC_t^i - CPC_{t-k}^i) + \sum_j mean(CPC_t^j - CPC_{t-k}^j) mean(CPC_t^i - CPC_{t-k}^i) + \sum_j cov(CPC_t^j - CPC_{t-k}^j, CPC_t^i - CPC_{t-k}^i)}{MSE(S_t - S_{t-k})} \quad j = 1(TY), \dots, 5(PA), j \neq i. \quad (12)$$

The difference between the two measures (11) and (12) is that the former ignores comovements, while the latter attributes half of the comovements to  $CPC_t^i$ . As discussed in the introduction, the CPCs may be dependent with each other. If they are correlated, the two measures show distinct results.

\*\*\*\*\*Figure 3 around here\*\*\*\*\*

Figures 3-1 and 3-2 show the results of a decomposition calculated with the measures (11) and (12). Both figures show that the longer is the horizon, the more largely the London/N.Y. segment contributes to the movement of yen/dollar rate. Specifically, for a short horizon, the Tokyo and the London segments contribute to changes in the yen/dollar rate, but this effect is dominated by the contribution of the London/N.Y. segment if the horizon increases beyond about 60 days (2 months). In other words, the largest part of the yen/dollar rate change stems from changes in the London/N.Y. segment at a long horizon. Another notable feature is that the CPC of the Pacific segment is the least persistent among five CPCs. There is little difference in ordering of CPCs between the measures (11) and (12). This suggests that taking into account the dependency of the CPCs does not lead to different results.

#### 4. Concluding Remarks

This paper examines the price informativeness of the intraday yen/dollar exchange rate and considers the relationship between trading volume and informational efficiency. The empirical evidence shows that the daily exchange rate has a cointegrating relationship with the cumulative price change of the overlapping business hours of the London and New York markets, but not with that of any other market segments. We also find that the cumulative price change of the London/N.Y. segment is the most persistent and the highest contributor to daily exchange rate fluctuations among the five

market segments in the medium- and long-run, while the CPC of the Pacific segment is the least persistent for any time horizon.

These results suggest that the greatest concentration of informed traders is in the London/N.Y. segment and the smallest concentration of informed traders is in the Pacific segment. Given that the intraday transactions are the highest in the London/N.Y. segment and the lowest in the Pacific segment, these findings are consistent with the asymmetric information view of Admati and Pfleiderer (1988) that prices are more informative when trading volume is heavier.



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**Table 1**

Table 1: Regional time zone of five segments

| Activity (GMT)     | Tokyo     | London     | London/N.Y. | N.Y.        | Pacific     |
|--------------------|-----------|------------|-------------|-------------|-------------|
| No DST             | 0:00-8:00 | 8:00-13:30 | 13:30-16:00 | 16:00-22:00 | 22:00-24:00 |
| Only Eur. DST      | 0:00-8:00 | 8:00-13:30 | 13:30-15:00 | 15:00-22:00 | 22:00-24:00 |
| Eur. and Amer. DST | 0:00-8:00 | 8:00-12:30 | 12:30-15:00 | 15:00-21:00 | 21:00-24:00 |
| Only Amer. DST     | 0:00-8:00 | 8:00-12:30 | 12:30-16:00 | 16:00-21:00 | 21:00-24:00 |

Note: ‘DST’ is daylight saving time. ‘Eur’ and ‘Amer’ refer to Europe and America. Since we assume that the London opening equal to the Tokyo closing, the London opening is fixed at 8:00am GMT.

**Table 2**

Table 2: Summary of statistics

| Variable         | N.o.b | Mean   | Per-hour variance | Variance ratio |
|------------------|-------|--------|-------------------|----------------|
| $\Delta P^{TKO}$ | 5126  | -0.007 | 0.024             | <b>2.385</b>   |
| $\Delta P^{LND}$ | 5126  | -0.001 | 0.031             | <b>1.877</b>   |
| $\Delta P^{L/N}$ | 5126  | -0.005 | 0.058             | -              |
| $\Delta P^{NY}$  | 5126  | 0.007  | 0.021             | <b>2.808</b>   |
| $\Delta P^{PAC}$ | 5126  | 0.001  | 0.025             | <b>2.352</b>   |

Note: N.o.b. is the number of observations in the sample. The price changes of each segment are divided by the root square of hours in the segment. “Variance ratio” denotes the variance ratio of the returns of N.Y. to those of each segment. Bold indicates a rejection of the null that the ratio is unity at the 1% significance level.

**Table 3**

Table 3: Unit root tests for CPC and the daily yen/dollar

The ADF unit root test for *level*

| Lag | Tokyo  | London | London/N.Y. | N.Y.   | Pacific | daily rate |
|-----|--------|--------|-------------|--------|---------|------------|
| 1   | -0.132 | -1.946 | -1.100      | -0.044 | -0.820  | -2.579 *   |
| 2   | -0.066 | -1.984 | -1.128      | -0.004 | -0.884  | -2.534     |
| 3   | -0.006 | -1.983 | -1.075      | 0.012  | -0.864  | -2.544     |
| 4   | 0.095  | -1.979 | -1.059      | 0.000  | -0.914  | -2.462     |
| 5   | 0.097  | -2.062 | -1.076      | 0.001  | -1.027  | -2.514     |

The ADF unit root test for *first-difference*

| Lag | Tokyo      | London     | London/N.Y. | N.Y.       | Pacific    | daily rate  |
|-----|------------|------------|-------------|------------|------------|-------------|
| 1   | -53.83 *** | -49.31 *** | -51.28 ***  | -52 ***    | -48 ***    | -51.302 *** |
| 2   | -44.02 *** | -40.73 *** | -42.53 ***  | -42.18 *** | -40.03 *** | -41.821 *** |
| 3   | -39.03 *** | -35.45 *** | -36.84 ***  | -36.12 *** | -34.31 *** | -37.274 *** |
| 4   | -34.23 *** | -30.4 ***  | -32.67 ***  | -32.51 *** | -29.16 *** | -32.065 *** |
| 5   | -30.92 *** | -28.26 *** | -28.98 ***  | -30.45 *** | -27.4 ***  | -29.848 *** |

Note: The reported numbers are ADF test statistics. The ADF test statistic is calculated with the assumption of whiteness of residuals (Hamilton (1994)). “Lag” in the first column denotes the lag length for the ADF test. The daily rate is the Tokyo opening rate. \*\*\* and \* indicate 1% and 10% significance, respectively. The critical values are -3.430 (1%), -2.860 (5%), and -2.570 (10%).

**Table 4**

Table 4: ADF cointegration test

| Lag | Tokyo  | London | London/N.Y. | N.Y.   | Pacific |
|-----|--------|--------|-------------|--------|---------|
| 1   | -2.363 | -2.304 | -2.883 **   | -2.559 | -2.531  |
| 2   | -2.367 | -2.283 | -2.871 **   | -2.553 | -2.528  |
| 3   | -2.337 | -2.249 | -2.850 *    | -2.524 | -2.504  |
| 4   | -2.286 | -2.194 | -2.751 *    | -2.464 | -2.442  |
| 5   | -2.360 | -2.232 | -2.844 *    | -2.523 | -2.484  |

Note: The reported numbers are ADF test statistics. “Lag” in the first column is the lag length for the ADF test. \*\* and \* indicate 5 and 10% significance, respectively. The critical values are -3.430 (1%), -2.860 (5%), and -2.570 (10%).

**Table 5**

Table 5: Testing for structural breaks

|             | Segment     | Break date | Minimum Phillips $Z_t^*$ test statistic |
|-------------|-------------|------------|-----------------------------------------|
| Model C     | Tokyo       | 1993/03/09 | -3.28                                   |
|             | London      | 1992/09/11 | -3.26                                   |
|             | London/N.Y. | 2004/06/10 | -3.57                                   |
|             | N.Y.        | 1993/03/09 | -3.51                                   |
|             | Pacific     | 2002/05/23 | -3.43                                   |
| Model C/S   | Tokyo       | 1993/01/27 | -3.30                                   |
|             | London      | 1993/01/14 | -3.22                                   |
|             | London/N.Y. | 1996/05/09 | -4.20                                   |
|             | N.Y.        | 1992/09/02 | -4.24                                   |
|             | Pacific     | 1995/06/29 | -4.24                                   |
| Model C/S/T | Tokyo       | 1995/10/20 | -3.99                                   |
|             | London      | 1996/11/01 | -3.68                                   |
|             | London/N.Y. | 1998/10/13 | -6.26 ***                               |
|             | N.Y.        | 1990/10/19 | -5.09                                   |
|             | Pacific     | 1995/08/14 | -4.17                                   |

Note: Models C, C/S and C/S/T assume endogenous changes in intercept, intercept/slope, and intercept/slope/trend, respectively. “Break date” refers to a possible structural break point in a cointegration vector. “The minimum Philips  $Z_t^*$  test statistic” corrects a bias due to a serial correlation of disturbance terms. \*\*\* indicates 1% significance. The critical values (Gregory and Hansen, 1996a, b) are

Model C: -5.13 (1%), -4.61 (5%), -4.34 (10%).

Model C/S: -5.47 (1%), -4.95 (5%), -4.68 (10%).

Model C/S/T: -6.02 (1%), -5.50 (5%), -5.24 (10%).

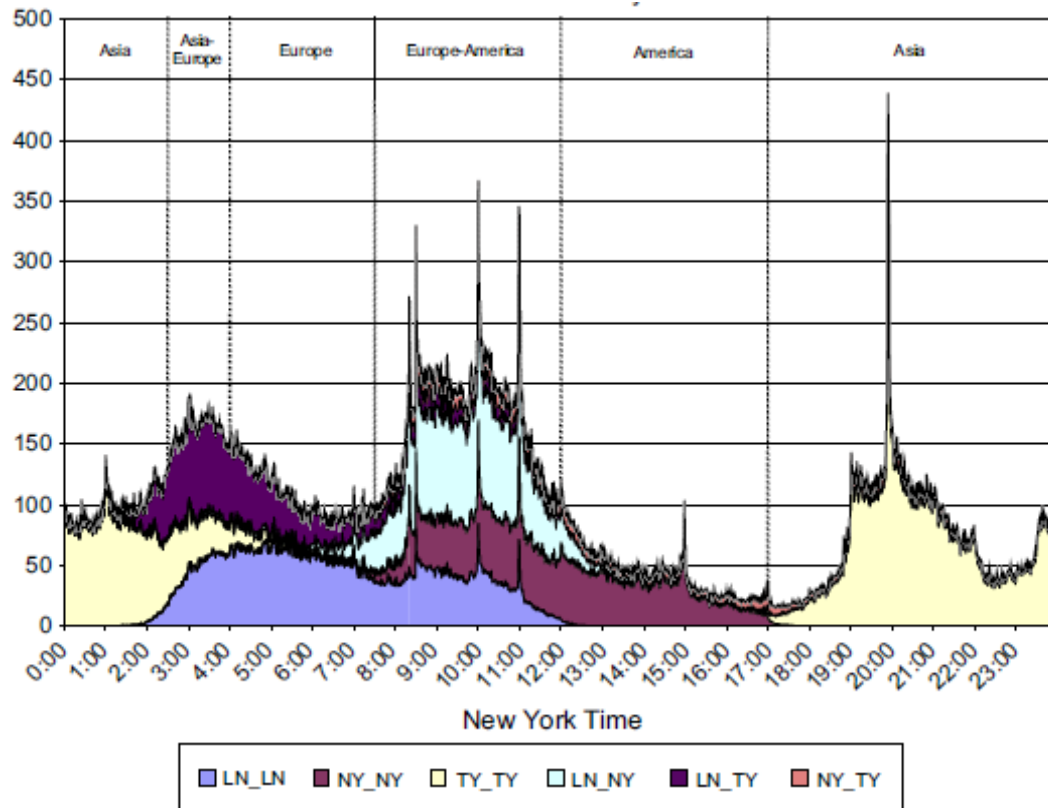
**Table 6**

Table 6: Variance ratio tests

| Lag | Statistics | Segment      |        |              |       |             |
|-----|------------|--------------|--------|--------------|-------|-------------|
|     |            | Tokyo        | London | London/N.Y.  | N.Y.  | Pacific     |
| 5   | VR         | 0.83         | 1.04   | 0.92         | 0.97  | 1.10        |
|     | L=M        | <b>-3.91</b> | 0.63   | <b>-2.20</b> | -0.74 | <b>2.62</b> |
|     | Rank       | <b>-3.50</b> | 0.82   | <b>-2.51</b> | -1.24 | <b>2.86</b> |
|     | Sign       | -0.32        | 1.41   | -0.84        | -0.10 | <b>3.21</b> |
| 10  | VR         | 0.78         | 1.08   | 0.91         | 0.92  | 1.20        |
|     | L=M        | <b>-3.42</b> | 0.88   | -1.62        | -1.15 | <b>3.19</b> |
|     | Rank       | <b>-2.79</b> | 0.88   | -1.62        | -1.51 | <b>3.76</b> |
|     | Sign       | 0.75         | 1.18   | 0.20         | 0.51  | <b>3.46</b> |
| 20  | VR         | 0.69         | 1.08   | 0.92         | 0.90  | 1.33        |
|     | L=M        | <b>-3.30</b> | 0.63   | -1.00        | -1.09 | <b>3.66</b> |
|     | Rank       | <b>-2.61</b> | 0.59   | -0.50        | -1.57 | <b>4.50</b> |
|     | Sign       | 1.58         | 1.55   | 1.50         | -0.03 | <b>4.11</b> |
| 30  | VR         | 0.65         | 1.02   | 0.95         | 0.85  | 1.46        |
|     | L=M        | <b>-3.01</b> | 0.15   | -0.50        | -1.31 | <b>4.11</b> |
|     | Rank       | <b>-2.32</b> | 0.42   | 0.37         | -1.90 | <b>4.98</b> |
|     | Sign       | 1.84         | 1.96   | <b>2.65</b>  | -0.37 | <b>4.92</b> |

Note: "VR" is the variance-ratio estimate. "L=M" is Lo and Mackinlay's (1988) heteroskedasticity-robust z-statistic. "Rank" and "Sign" are Wright's (2000) rank- ( $R_2$ ) and signed-based ( $S_1$ ) statistics, respectively. Bold denotes statistical significance at the 5% level. The critical value is based on Table 1 in Wright (2000)

Figure 1: Intraday trading volume pattern in the yen/dollar foreign exchange market



Note: This figure shows regional trading volumes for Tokyo, London, and N.Y. All trading volumes are indexed to the average overall trading volume where 100 equal the average overall trading volume per 1-min period over the whole sample. Regional mnemonics are LN-LN (London and London), NY-NY (N.Y. and N.Y.), TY-TY (Tokyo and Tokyo), LN-NY (London and N.Y.), LN-TY (London and Tokyo), and NY-TY (N.Y. and Tokyo).

Citation: Cai et al. (2007)



Figure 2-1: The Tokyo CPC (solid line) and the daily yen/dollar rate (dotted line)

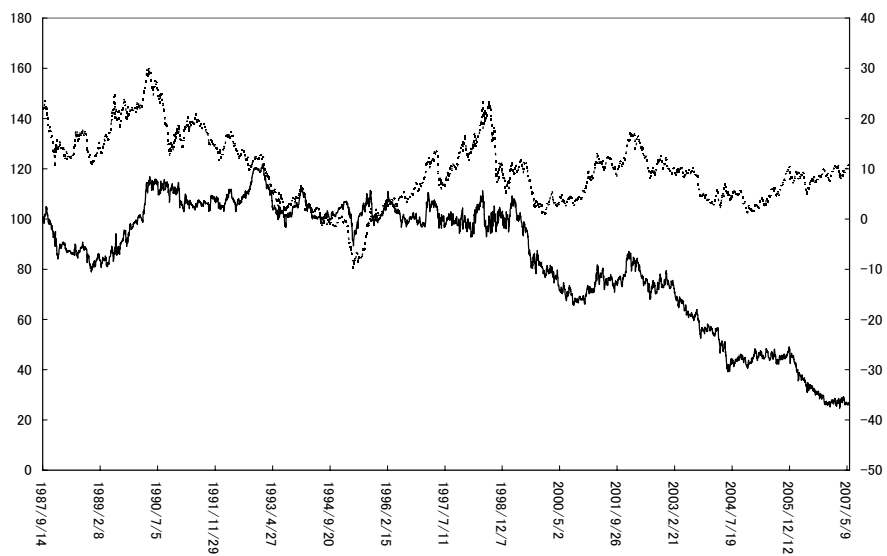


Figure 2-2: The London CPC (solid line) and the daily yen/dollar rate (dotted line)

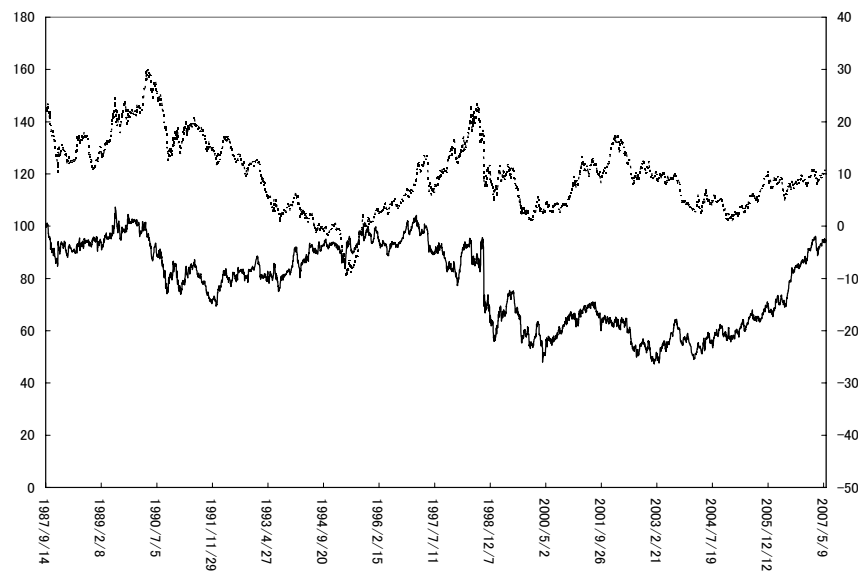


Figure 2-3: The London/N.Y. CPC (solid line) and the daily yen/dollar rate (dotted line)

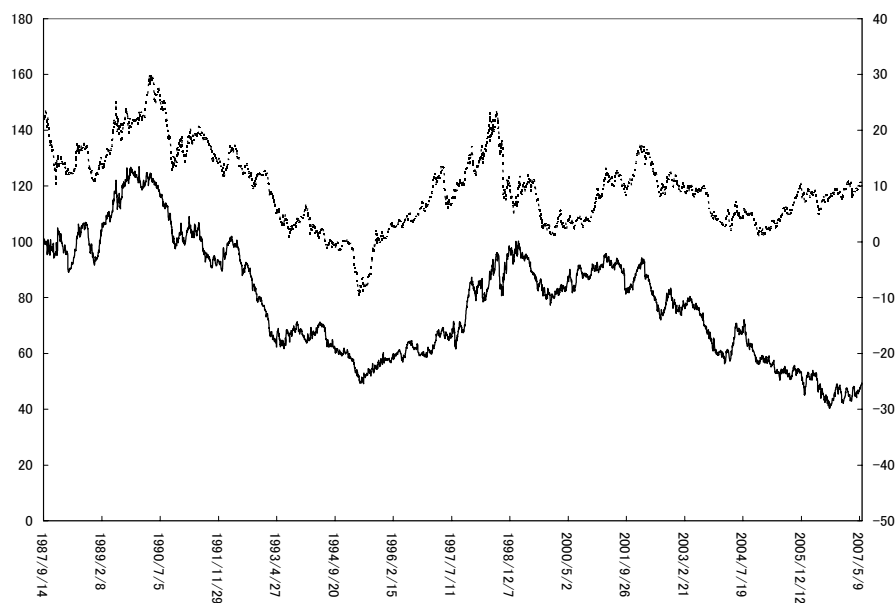
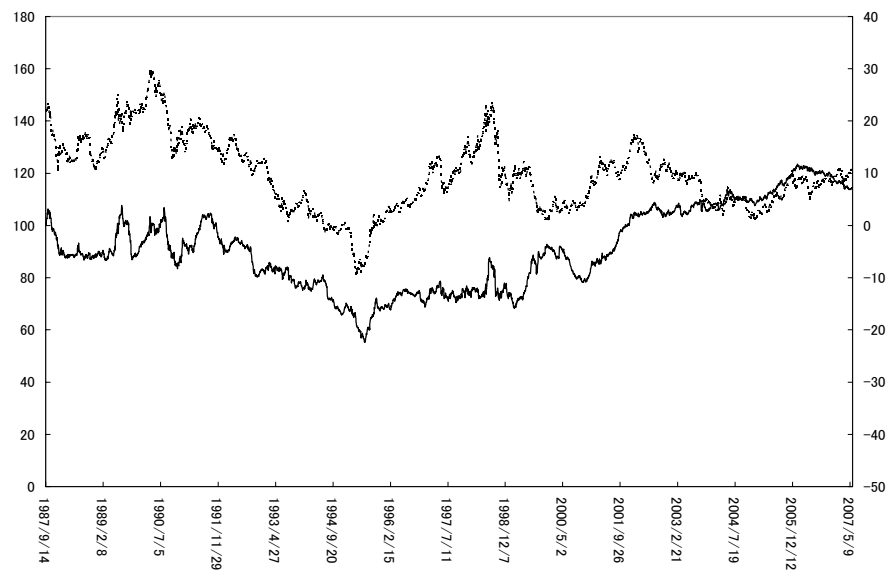


Figure 2-4: The N.Y. CPC (solid line) and the daily yen/dollar rate (dotted line)



Figure 2-5: The Pacific CPC (solid line) and the daily yen/dollar rate (dotted line)



Note: Each CPC at  $t$  corresponds to the Tokyo opening rate at  $t+1$ . The sample period is September 14, 1987 to May 31, 2007.

**Figure 3**

Figure 3-1: MSE Decomposition [Measure (11)]

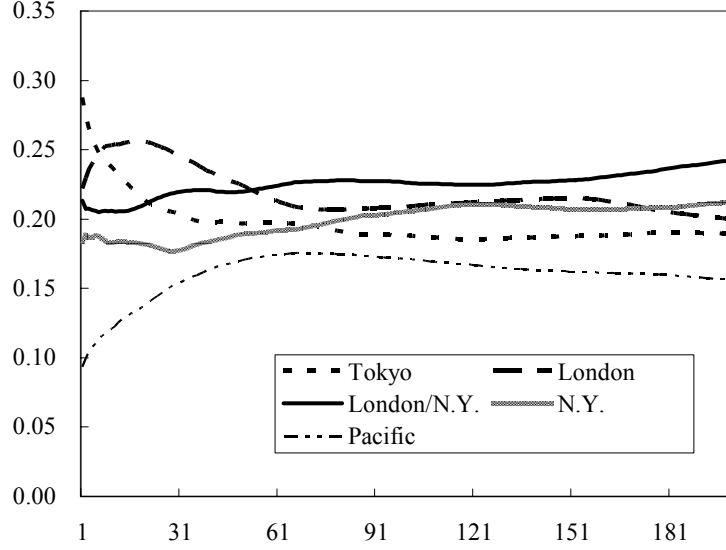
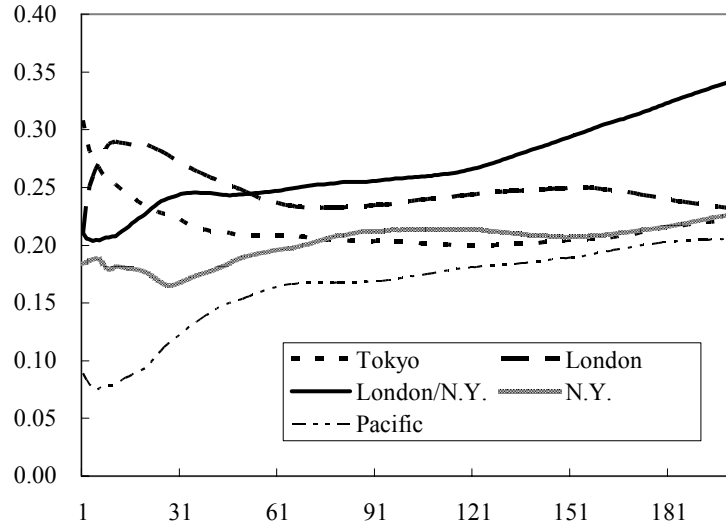


Figure 3-2: MSE Decomposition [Measure (12)]



Note: These two figures are calculated with the measures (11) and (12).

$$\frac{MSE(CPC_t^i - CPC_{t-k}^i)}{\sum_i MSE(CPC_t^i - CPC_{t-k}^i)} \quad i = 1(TY), \dots, 5(PA), \quad (11)$$

$$\frac{MSE(CPC_t^i - CPC_{t-k}^i) + \sum_j mean(CPC_t^j - CPC_{t-k}^j) mean(CPC_t^i - CPC_{t-k}^i) + \sum_j cov(CPC_t^j - CPC_{t-k}^j, CPC_t^i - CPC_{t-k}^i)}{MSE(S_t - S_{t-k})}$$

$$j = 1(TY), \dots, 5(PA), j \neq i. \quad (12)$$