JOURNAL OF APPLIED ECONOMETRICS, VOL. 10, 447–467 (1995)

# JAPAN'S FINANCIAL DEREGULATION AND LINKAGE OF THE GENSAKI AND EUROYEN DEPOSIT MARKETS

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SUMMARY

This paper examines whether Japan's financial deregulation weakened the linkage between the Euroyen and Gensaki markets. It defines and analyses the linkage of the two markets in terms of the persistence and predictability of yield spreads, as well as cross-market causality. By using Johansen tests and an error correction model, this paper shows first, cross-market causality before and after deregulation; and second, one common trend in interest rates, substantial Gensaki–Euroyen yield spreads, and a slower response of the one-month Gensaki–Euroyen yield spread to a shock after deregulation.

## 1. INTRODUCTION

One of the phenomena in money markets is that assets with similar risk characteristics can trade at different prices (e.g. Amihud and Mendelson, 1991; Stigum 1990). The existence of substantial interest rate differentials is often caused by risk premiums or market frictions such as differences in transaction costs, legal barriers, and liquidity in the markets. (See Boudoukh and Whitelaw, 1993 and Delgado and Dumas 1993.) One typical example is the interest rate differential between on- and offshore money market instruments. In the past two decades, industrial countries have progressively removed obstacles to international capital flows and dismantled regulatory barriers in onshore markets. This ongoing financial deregulation can reduce the market frictions, increase international capital mobility, and narrow the interest rate differentials between on- and offshore markets. The goal of this paper is to explore this issue empirically by defining the linkage of two money markets which trade assets with similar risk characteristics and by examining the effect of Japanese financial deregulation on the linkage between on- and offshore money markets.

This paper defines the linkage of the two money markets in terms of three conditions: (1) the stationarity of interest rate differentials; (2) cross-market causality in interest rates; (3) the adjustment speed of interest rate differentials. The first condition is to rule out apparent long-term profit opportunities. If two markets are linked, interest rates should not drift apart in the long run when a shock occurs. If two interest rates drift apart, then the interest rate differential will eventually exceed any finite transaction costs and the arbitrage can take place by lending high and borrowing low.[1] Therefore, the stationarity of the interest rate differentials ensures that there is no apparent long-term profit opportunity. Applying the

[1] If the interest rate differential $x_t$ follows a random walk with a drift $\mu$, then the expected interest rate differential at period $t$ after the occurrence of the shock at period 0 will be $\mu t + x_0$. The expected interest rate differential will eventually exceed any finite transaction cost and the arbitrage can be taken at that time.

CCC 0883–7252/95/040447–21

*Received April 1993*
*Revised October 1994*

concept of cointegration in Engle and Granger (1987), these interest rates are cointegrated and have common trends.

The second condition is frequently used in the literature to examine the lead and lag relation of asset prices. Such a relation is often interpreted as the transmission of information flow across markets. For example, see Chan, Chan, and Karolyi (1991), Engle, Ito, and Lin (1990), and Lin, Engle, and Ito (1994). These studies asked whether returns (or volatility) from one market can provide useful information in explaining returns (or volatility) in another market. If two markets are linked, a bi-directional, cross-market causality will be exhibited.

The third condition is to guarantee that there is no short-run profit opportunity so that investors cannot exploit the profit by using simple trading strategies based on past information. However, even if a market is efficient, the predictability of interest rate differentials may exist due to different risk premia and market frictions. Unpredictability is too restrictive in most empirical studies. An alternative way to measure the market linkage is to examine the relative adjustment speed of these markets in response to a shock in order to measure how fast the interest rate differentials die out.

The above three conditions are analysed in the framework of cointegration and a vector error-correction (ECM) model. My empirical interest will be centred on the effect of deregulation on the linkage of the Euroyen and Gensaki markets—a repurchase agreement market for government bonds.[2] Japanese deregulation of onshore certificates of deposit (CD) markets began in 1979. Before such deregulation, the Gensaki market was actively used as an open market to move funds between suppliers and demanders in the short-term money market. Deregulation has decreased the depth of the Gensaki markets (e.g. Takagi, 1987) due to the emergence of new assets and the existence of security transfer taxes. This decline in market depth can weaken the linkage between the Gensaki and (short-term) Euroyen deposit markets, an offshore market free of onshore market regulations, and challenge the role of the Gensaki market as a representative short-term money market.

Studies by Campbell and Hamao (1992) and Singleton (1990) have highlighted the importance of Japan's financial deregulation on pricing money market instruments and government bonds. In contrast to this article, these studies focus only on the domestic onshore market. A related paper by Bonser-Neal and Roley (1993) examined a bivariate cointegration relation between onshore market interest rates (e.g. Japanese Gensaki or CD) and offshore market interest rates (e.g. Euroyen deposit or Eurodollar deposit rates) over every two-period horizon from 1980 to 1990. This paper adopts a different definition of market linkage from that in Bonser-Neal and Roley (1993). While a cointegration relation is required for both models, in contrast to the equal variances of two interest rates used in their paper, this paper examines the market linkage based on cross-market causality and the predictability of Gensaki–Euroyen yield spreads.

The weekly series of Gensaki and Euroyen deposit rates from December 1980 to September 1989 are analysed. To carry out the analyses, I conduct the following three econometric procedures. First, the Johansen test (1991) with a structural break in short-run dynamics is used to examine the number of common trends in these interest rates series. This test is based on the rank of canonical correlations between the levels and the first differences of data after correcting for any short-run dynamics and allowing for a break in the short-run dynamics. A joint test for a known structural break in the cointegration matrix and its rank is also conducted within the Johansen test framework. Second, the error correction model is estimated and

[2]The repurchase agreement (PR) is a short-term investment in which sellers (usually bonds dealers in Japan) borrow from an investor (usually a corporation) using one security as collateral and agree to repurchase it at an agreed price on a stated day.

Granger causality for cross-market information is defined in the context of the error correction model. Third, applying methods similar to those in Lütkepohl and Reimers (1992) and Warne (1991), the impulse response functions of these spreads and their standard errors are derived to inspect the speed of the market adjustment to a shock. Wald tests for the unpredictability of Gensaki–Euroyen yield spreads (interest rate differentials) and for the speed of yield spreads in response to a shock are constructed and performed.

This paper proceeds as follows. Section 2 sets out the econometric methods used in this analysis. Section 3 provides a data summary for interest rates and yield spreads and the empirical results for the number of common trends. Section 4 estimates an error-correction model and derives the impulse response function for yield spreads. Section 5 concludes.

## 2. AN ERROR-CORRECTION MODEL FOR THE TERM STRUCTURE OF INTEREST RATES

### 2.1. Common Trends in Yield Curves

The log linear version of the term structure of interest rates asserts that long-term interest rates are a weighted average of the expected future short-term interest rates and the expected excess holding period returns (rolling term premium). Let $t$ denote the $t$th week and one month contain approximately four weeks. Define $r_{1,t}$ as the yield on the one-month pure discount bond (also denoted as one-month Gensaki rates) and $r_{n,t}$ as the yield on the $n$-month pure discount bond. Then $r_{n,t}$ obeys

$$r_{n,t} = (1/n) \sum_{k=0,n-1} \mathrm{E}_t(r_{1,t+4k}) + \phi_{n,t} \tag{1}$$

where $\phi_{n,t}$ is the rolling term premium.

Equation (1) shows that long-term interest rates are driven by two components—expectations for future short-term interest rates and the rolling term premium. I begin by assuming that one-month Gensaki rates can be represented as a process of integrated of order one, $\tau_t$, and a stationary process, $w_t$:

$$r_{1,t} = \tau_t + w_t \tag{2}$$

Using equations (1) and (2), I obtain:

$$r_{n,t} = \tau_t + w_t + (1/n) \sum_{k=1,n-1} \left\{ \left[ \sum_{j=1,4k} \mathrm{E}_t(\Delta\tau_{t+j}) \right] + \mathrm{E}_t(w_{t+4k}) \right\} + \phi_{n,t} \tag{3}$$

where $\Delta$ is the first difference operator. Equations (2) and (3) imply that if $r_{n,t}$ for $n = 1,3$ is non-stationary and the rolling term premium is stationary, then the entire spectrum of interest rates has one common trend. Since the Euroyen deposit market offers the yen-denominated money market instruments which are substitutes for Gensaki transactions, the absence of apparent long-term profit opportunities between the two markets implies that the one-month Euroyen deposit rate follows a trend component $\tau_t$, common to that of the one-month Gensaki rate, and a transitory component $w_t^*$:

$$r_{1,t}^* = \tau_t + w_t^* \tag{4}$$

Equations (2)–(4) indicate that there is one common trend in the term structure of the Gensaki and Euroyen deposit rates.

### 2.2. Johansen Test for the Number of Common Trends in the Presence of a Known Structural Break

My primary interest is in examining the effect of Japanese financial deregulation on Gensaki and Euroyen interest rates, so I assume that the Gensaki and Euroyen deposit rates have the following vector autoregressive (VAR) representation with a structural break in the parameters of the process:

$$y_t = \mu_1 + \sum_{k=1,p} A_{k1} y_{t-k} + u_t \qquad \text{for } t = 1, ..., \tau \tag{5a}$$

and

$$y_t = \mu_2 + \sum_{k=1,p} A_{k2} y_{t-k} + u_t \qquad \text{for } t = \tau + 1, ..., T \tag{5b}$$

where $y_t = (r_{1,t}, r_{3,t}, r^*_{1,t}, r^*_{3,t})'$; $r_{m,t}$ and $r^*_{m,t}$ are the $m$-month Gensaki and Euroyen deposit rates for $m = 1, 3$; $\tau$ is a break point; and $T$ is the number of observations. To perform hypothesis testing, it is convenient to rewrite equations (5a) and (5b) as the ECM form:

$$\Delta y_t = \mu_i + \sum_{k=1,p-1} \Gamma_{ki} \Delta y_{t-k} + \Pi_i y_{t-1} + u_t \qquad \text{for period } i = 1, 2 \tag{6}$$

where $\Gamma_{ki} = -\Sigma_{j=k+1,p} A_{ji}$, $\Pi_i = \Sigma_{j=1,p} A_{ji} - I$. Johansen (1991) and Johansen and Juselius (1990) examine a cointegration relation based on the hypothesis of the rank of $\Pi_i$. If the rank of $\Pi_i$ is equal to $q$ and less than the number of variables, $n$. $\Pi_i$ can be expressed as

$$\Pi_i = \alpha_i \beta_i' \qquad \text{for period } i = 1, 2$$

where $\alpha_i$ and $\beta_i$ are $n \times q$ matrices, $\alpha_i$ is called the adjustment matrix, and $\beta_i$ is called the cointegration matrix. Johansen (1991) and Johansen and Juselius (1990) have provided an excellent exposition of conducting the test for the rank of the cointegration matrix under stable long- and short-run dynamics. The test procedure with a structural break has been developed by Hansen and Johansen (1993), who emphasize the role of the stability of the short-run parameters in testing the stability of the long-run parameters and the rank of the cointegration matrix. If the short-run parameters are stable, it is better to fix the short-run parameters at the full-sample values to decrease the variance of the estimated parameters. If the short-run parameters are unstable, it is better to re-estimate short-run parameters to obtain the correct size of the Johansen tests.[3] This paper considers the test under unstable short-term parameters. To avoid repetition, this section gives a brief description of the procedures for conducting Johansen tests in the presence of unstable short-run dynamics.

In the first step, I assume that the constant vector $\mu$ appears in the cointegration matrix when I test for the rank of the cointegration matrix. This maintained assumption will be tested later. To account for a structural break in the short-run dynamics, under the maintained assumption that $\mu_i = \alpha_i \delta_i$ and $\Gamma_{k1} \neq \Gamma_{k2}$ for $k = 1, \ldots, p-1$, the null hypothesis for each subperiod is set as

$$\mathrm{H}_{q,u}\text{: } \mathrm{rank}(\Pi_i) = q_i \qquad \text{for period i} = 1, 2 \tag{7}$$

The test proposed by Johansen and Juselius (1990) can be performed for each subperiod. Let $y^*_t = (r_{1,t}, r_{3,t}, r^*_{1,t}, r^*_{3,t}, 1)'$ and $Z_{t-1} = (\Delta y'_{t-1}, \ldots, \Delta y'_{t-p+1})'$. By assuming that the rank of the

[3] Cheung and Lai (1992) showed in a Monte Carlo study that the size of the Johansen test is distorted if the error terms have a moving average process. Misspecifications of the short-run parameters may induce serial correlations in residuals in the first-step regression. The size of the Johansen tests may be affected.

cointegration matrix and the long-run relation are stable for the two subperiods, the test for the whole sample period can be conducted first, by correcting for short-run dynamics by regressing $y_t^*$ and $\Delta y_t$ on $Z_{t-1}$ and obtaining their residuals $\varepsilon_{1,t}$ and $\varepsilon_{0,t}$, respectively, and second, by combining residuals from the two periods and using the reduced rank regression of $\varepsilon_{0,t}$ on $\varepsilon_{1,t}$. Under the maintained assumption that $\Pi_1 = \Pi_2 = \Pi$, $\mu_i = \alpha\delta_i$, $\Gamma_{j1} \neq \Gamma_{j2}$ for j = $1,\ldots,p-1$, and period $i = 1, 2$, the null hypothesis is set to be

$$\mathrm{H}_{q,s}\colon \operatorname{rank}(\Pi) = q \tag{8}$$

In the second step, some tests of model specifications will be conducted to give assurance about model specified above and to shed light on the first condition for market linkage. The first test examines whether the constant vector $\mu$ appears only in the cointegration matrix. A constant vector in the ECM in equation (6) generally gives rise to a linear deterministic trend in $y_t$. As discussed in Johansen (1993), if the constant vector $\mu$ lies in the space spanned by $\alpha$, the constant vector appears only in the cointegration matrix and the variables in levels $y_t$ do not have a linear deterministic trend. This test is performed under the maintained assumption that short-run dynamics are unstable and the rank of the cointegration matrix is known; the null hypothesis is

$$\mathrm{H}_{\mu}\colon \mu_1 = \alpha_i \delta_t \text{ for period i} = 1, 2 \tag{9}$$

where $\delta_i$ is a $q \times 1$ vector. After the conclusion of the rank of the cointegration matrix is drawn, the second test examines the stability of long-run parameters and the rank of the cointegration matrix under unstable short-run parameters, which can be formulated as

$$\mathrm{H}_{s,u}\colon \Pi_1 = \Pi_2 \tag{10}$$

The likelihood ratio test (LR) for a structural break in the long-run parameters and the rank of the cointegration matrix is conducted by computing the ratio of the value of the log likelihood function for the whole period to the sum of the values for the two subperiods. Under the null hypothesis that the rank of the cointegration matrix for the whole period is at least greater than any for the two subperiods, Quinto (1993) shows that this test has a $\chi^2$ distribution. The third test is the restriction on the loading of the common trend in the interest rate processes (i.e. the same response of interest rates to the common trend). This hypothesis implies that any yield spreads (differences between two interest rate series) form a cointegrating vector. For example, the vector, (0, 0, 1, $-1$), is a cointegrating vector. Thus, it imposes constraints on the cointegration matrix $\beta$, that is, $\beta = H\varphi$, where $H$ is a pre-specified matrix.

### 2.3. Granger Causality and Impulse Response Function

After concluding that an error-correction model is appropriate for the data, I set up the model to characterize the short-run dynamics of interest rates for each subperiod. In the following analysis, I drop subscript $i$, for period $i = 1, 2$, to simplify the notation. Analysis can be applied for the two subperiods separately.

There are many possible yield spreads which might serve as cointegrating vectors for interest rates. I use the spreads that are most able either to account for the effect of the deregulation on the linkage of Japanese onshore and offshore markets or to predict future short-term interest rates. Specifically, the spread ($s_{1,t}$) between one- and three-month Gensaki rates, the spread ($s_{2,t}$) between one-month Gensaki and Euroyen deposit rates, and the spread ($s_{3,t}$) between three-month Gensaki and Euroyen deposit rates are chosen for these two purposes. By assuming that the order of the autoregressive processes in levels equals two, the error-correction representation for the interest rate vector $y_t$ can be written as

$$\Delta y_t = \Gamma \Delta y_{t-1} + \alpha(s_{t-1} - \delta) + u_t \tag{11}$$

where $s_t = (s_{1,t}, s_{2,t}, s_{3,t})'$.

### 2.3.1. Granger Causality

To verify the second condition for the linkage of the two markets, I examine whether the cross-market information is useful in predicting interest rates in one market by testing for no Granger-causality between two market interest rates in the ECM. Let $r_t = (r_{1,t}, r_{3,t})'$, $r_t^* = (r_{1,t}^*, r_{3,t}^*)'$, and $s_t^* = (s_{2,t}, s_{3,t})'$. The decompositions of $y_t$ and $s_t$ according to $r_t$ and $r_t^*$ can be written as $(r_t', r_t^*)'$ and $(s_{1,t}', s_t^{*\prime})'$, respectively. Let also the decompositions of $\Gamma$ and $\alpha$ be conformable to the decomposition of $y_t$ and $s_t$.

Under the null hypothesis of no Granger-causality from $r$ to $r^*$, the lower off-diagonal submatrices of $A_k$ for all $k$ in the VAR for levels (i.e. equations (5a) and (5b)) are zero. Equivalently, as shown in Mosconi and Giannini (1992), the Granger-causality test in the ECM on equation (6) examines the following two hypotheses. The null hypothesis that $r$ does not Granger cause $r^*$ can be formulated as

$$\mathrm{H}_{G,r}\colon \Gamma_{ij} = 0 \text{ for all } i = 3,4 \quad \text{and} \quad j = 1,2, \quad \text{and} \quad \alpha_{ij} = 0 \text{ for all } i = 3,4 \quad \text{and} \quad j = 1,2,3 \tag{12}$$

where $\Gamma_{ij}$ is the $ij$th element of $\Gamma$ and $\alpha_{ij}$ is the $ij$th element of $\alpha$. Similarly, the null hypothesis that $r^*$ does not Granger-cause $r$ can be formulated as

$$\mathrm{H}_{G,r^*}\colon \Gamma_{ij} = 0 \text{ for all } i = 1,2 \quad \text{and} \quad j = 3,4 \quad \text{and} \quad \alpha_{ij} = 0 \text{ for all } i = 1,2 \quad \text{and} \quad j = 2,3 \tag{13}$$

### 2.3.2. Impulse Response Function

An impulse response function traces the dynamic effect of a shock on economic or financial variables. Lutkepohl and Reimers (1992) derived such a response curve in a cointegrating VAR system. They treat the impulse response function as a nonlinear function of parameters in the ECM and construct analytical derivatives. The response of variables in levels to orthogonal innovations is obtained by the approximation theorem for a nonlinear random function. The following derivation of the impulse response function is similar to that used in Lutkepohl and Reimers (1992) and Warne (1991). Note that

$$s_t = \beta'\alpha\delta + \beta'\Gamma\Delta y_{t-1} + (\beta'\alpha + I_3)s_{t-1} + \beta' u_t$$

where $I_3$ is a $3 \times 3$ identity matrix. By letting $x_t = (\Delta y_t', s_t')'$, we can form $x_t$, as a constrained vector autoregressive process:

$$x_t = \gamma + A x_{t-1} + G u_t \tag{14}$$

where

$$A = \begin{bmatrix} \Gamma & \alpha \\ \beta'\Gamma & I_3 + \beta'\alpha \end{bmatrix} \qquad G = \begin{bmatrix} I_3 \\ \beta \end{bmatrix} \qquad \text{and} \qquad \gamma = G\alpha\delta$$

The impulse response function of $x_t$ to one unit shock of $u_t$ for the $i$th step ahead can be obtained as

$$R_k = A^k G$$

Standard errors for $R_k$ can be calculated by $D_k \Omega D_k'$ where $D_k$ is the derivative of $R_k$ with respect to the parameter vector $\theta$, where $\theta = (\mathrm{vec}(\Gamma)', \mathrm{vec}(\alpha)')'$; $\Omega$ is the covariance matrix of parameters; and vec is an operator which transforms a matrix into a vector by stacking the columns of the matrix one underneath the other. Using the results from Magnus and Neudecker (1988), the analytical value of $D_k$ is equal to $\Sigma_{j=0,k-1} R'_{k-1-j} \otimes R_j$, where $\otimes$ is a Kronecker product.[4]

Another issue regarding the dynamic response of Gensaki–Euroyen yield spreads is whether the Gensaki and Euroyen deposit rates move synchronously so that their yield spreads become unpredictable. From equations (1) and (4), this condition implies that $w_t - w_t^* = \varepsilon_t$ or $w_t^* - w_t = \varepsilon_t$, where $\varepsilon_t$ is a white noise. This hypothesis that the one-month Gensaki–Euroyen yield spread is unpredictable in effect imposes restrictions on coefficients in equations of $\Delta r_{1,t}$, and $\Delta r^*_{1,t}$, that is, $h'\Gamma = 0$ and $h'\alpha = -1$, where $h' = [1, 0, -1, 0]$.

For a longer maturity, the Gensaki–Euroyen yield spreads will be affected by the movements of short-term interest rates and term premiums in both markets. The conditions for the unpredictability of the Gensaki–Euroyen yield spreads hinge on (1) the unpredictability of the one-month Gensaki–Euroyen yield spreads and (2) the synchronous movements of the term premium. Whether the dynamic effect of a shock on yield spreads with a longer maturity dies out more quickly after deregulation depends on these two forces.

## 3. EMPERICAL EVIDENCE ON THE COMMON TRENDS IN GENSAKI AND EUROYEN DEPOSIT RATES

This paper studies weekly series of one- and three-month Gensaki and Euro yen deposit rates from December 1980 to September 1989. The first week of December 1980 is chosen as a starting point because before that date, the capital control in Japan resulted in substantial Gensaki–Euroyen yield spreads. These yield spreads reflected the political risk (e.g. Ito, 1986). The data are obtained from the Japan Securities Dealers Association. The Euroyen deposit rates for one- and three-month maturities are obtained from the Bank for International Settlements.

The Gensaki market, an open and deregulated money market, played an important role in the short-term money market in the 1970s (e.g. Ito, 1992). Major buyers (business corporations) used the market as a vehicle for the efficient use of their surplus funds while major sellers (securities firms) considered the market as the way to finance their inventories of government bonds in the short run. Why did the depth of the Gensaki market decline during the evolution of Japan's financial deregulation? Several factors may have contributed to this phenomenon. For instance, the emergence of substitutes in the money market provided an alternative short-term

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[4] Using Table 7 (page 184) from Magnus and Neudecker (1988) and letting $d$ be a derivative operator and $R_k$ be $A^k G$,

$$d\,\mathrm{vec}(R_k) = \sum_{j=0,k-1} G' A^{k-1-j\prime} \otimes A^j d\,\mathrm{vec}(A)$$

Since $A = G[\Gamma \alpha] + J$, where

$$J = \begin{bmatrix} 0 & 0 \\ 0 & I_{p-1} \end{bmatrix}, d\,\mathrm{vec}(A) = (I \otimes G) d\,\mathrm{vec}(\theta)$$

The above two equations yield $D_k$.

financing instrument.[5] Liquidity in the secondary bond market increased when banks were allowed to trade on the secondary bond market (after June 1984) and bid–ask spreads were reduced in the secondary bond markets.[6] Also, the security transfer taxes and an increase in the volatility of ten-year government bonds might have changed market participants' assessments of taxes and risks in Gensaki.[7] According to reports in *Securities Markets in Japan* published by the Japan Securities Dealers Association, corporations' purchase of Gensaki accounted for 50–70% of total purchases before deregulation, but this declined to about 15% of total purchases in 1989. These factors led to changes in corporations' demand for the Gensaki and the security firms' inventory financing, resulting in a drop in Gensaki trading relative to CD or government bond trading.

Finding a single break point is essential and important in knowing the impact of financial deregulation on the linkage of Gensaki and Euroyen deposit markets. In this paper, I simply use the guidance of the mid-point criterion and economic events to infer a possible break point. A Monte Carlo simulation by Andrews, Lee, and Ploberger (1992) has shown that if a break point takes place in the beginning or at the end of the sample, the power of the test using the mid-point is much lower than that of the recursive Chow-type tests; if not, the power of the test using the mid-point is comparable to that of the recursive Chow-type tests. Hence, the test assuming a known break point is conservative to detect a break in the interest rate processes. With these considerations in mind, I chose the first week of August in 1985 as a break point (241st observation out of 452). This choice was based upon the change in monetary policy by the Bank of Japan after the Plaza Accord and the introduction of new money market assets in early 1985. This break point is also close to that chosen by Campbell and Hamao (1992), so I am able to compare my results with theirs.

### 3.1. Data Summary

Summary statistics for the data are given in Table 1. The second and third columns of Table I show the mean and standard deviations of interest rates and yield spreads for the first, the second, and the whole sample periods.[8] A notable and interesting phenomenon from panel A is a decrease in the sample mean of Gensaki–Euroyen yield spreads after August 1985. To get a clear picture of the fluctuations in the Gensaki–Euroyen yield spreads, Figures 1 and 2 plot the series of one- and three-month Gensaki–Euroyen yield spreads in the 1980s. These figures show that the variations in Gensaki–Euroyen yield spreads tend to narrow from 1980 to 1985 and that spreads tend to decline after the last quarter of 1985.

To check the stationarity of the data, the Dickey–Fuller (DF) (1979) tests for a unit root of interest rates and yield spreads are presented in the seventh and eighth columns for the first, the second, and the whole periods. A constant term, four lags of the first difference of the data

[5] For example, new assets such as bankers' acceptances and commercial papers were introduced in June 1985 and November 1987, respectively. This gives firms alternative ways to finance a short-term shortage of funds. Relaxation of the issuance, size, and maturity of certificates of deposit, and deregulation of the minimum amount of money market certificates and interest rates of large time deposits made these assets more attractive than the Gensaki because the securities transfer tax is levied on the Gensaki. Corporations have become a major buyer of these assets. (See Ito, 1992.)

[6] According to Takagi (1987), this deregulation has reduced transaction costs from 0.13% of annualized bond yields in 1980 to below 0.02% after 1985. However, the bid and ask prices for the Gensaki are not available.

[7] Using a Wald test, the standard deviations of weekly changes in 10-year government bond yields were found to increase significantly from 0.123 to 0.219 after August 1985.

[8] The standard errors are obtain by using the Generalized Method of Moments (Hansen, 1982) and the Newey and West (1987) autocorrelation-heteroskedasticity consistent covariance matrix with Bartlett kernels. The bandwidth in the covariance matrix is selected based on the Newey and West (1992) data-dependent procedure.

series, and one lag of the data series in levels are included as regressors. The estimated coefficient of one lag of the data series in levels (denoted as $a$) is also reported in the last column of Table I. In the first period, the unit root hypothesis is rejected for all four interest rates and yield spreads by using the DF(4) test statistics; the estimate of $a$ is quite close to zero, but significant from zero for all four interest rates and yield spreads. In the second period, the

Table I. Data summary

| | Mean | Std | $\rho(1)$ | $\rho(2)$ | $\rho(3)$ | DF(4) | $a$ |
|---|---|---|---|---|---|---|---|
| A. Before deregulation (period 1) | | | | | | | |
| $r_1$ | 6·610 | 0·663 | 0·805 | 0·550 | 0·306 | −4·229* | −0·063 |
| $r_3$ | 6·755 | 0·655 | 0·811 | 0·559 | 0·329 | −4·909* | −0·041 |
| $r_1^*$ | 6·644 | 0·678 | 0·585 | 0·347 | 0·215 | −4·202* | −0·110 |
| $r_3^*$ | 6·770 | 0·680 | 0·648 | 0·400 | 0·227 | −5·083* | −0·087 |
| $r_3-r_1$ | 0·145 | 0·161 | 0·846 | 0·621 | 0·424 | −6·356* | −0·314 |
| $r_3^*-r_1^*$ | 0·126 | 0·222 | 0·755 | 0·632 | 0·483 | −4·832* | −0·258 |
| $r_1-r_1^*$ | −0·034 | 0·214 | 0·028 | 0·179 | 0·031 | −6·899* | −0·867 |
| $r_3-r_3^*$ | −0·016 | 0·185 | 0·115 | 0·107 | −0·011 | −5·788* | −0·502 |
| B. After deregulation (period 2) | | | | | | | |
| $r_1$ | 4·505 | 0·946 | 0·961 | 0·896 | 0·826 | −1·961 | −0·013 |
| $r_3$ | 4·536 | 0·928 | 0·963 | 0·902 | 0·833 | −2·061 | −0·013 |
| $r_1^*$ | 4·804 | 0·990 | 0·849 | 0·829 | 0·724 | −2·038 | −0·035 |
| $r_3^*$ | 4·831 | 0·893 | 0·909 | 0·856 | 0·748 | −1·959 | −0·025 |
| $r_3-r_1$ | 0·031 | 0·069 | 0·850 | 0·634 | 0·414 | −4·611* | −0·191 |
| $r_3^*-r_1^*$ | 0·027 | 0·241 | 0·790 | 0·500 | 0·290 | −5·162* | −0·291 |
| $r_1-r_1^*$ | −0·299 | 0·270 | 0·608 | 0·445 | 0·327 | −4·712* | −0·299 |
| $r_3-r_3^*$ | −0·295 | 0·266 | 0·731 | 0·669 | 0·579 | −2·837 | −0·104 |
| C. Whole period | | | | | | | |
| $r_1$ | 5·623 | 1·326 | 0·982 | 0·961 | 0·939 | −4·084* | −0·031 |
| $r_3$ | 5·714 | 1·363 | 0·986 | 0·969 | 0·952 | −4·445* | −0·024 |
| $r_1^*$ | 5·781 | 1·244 | 0·959 | 0·932 | 0·905 | −4·057* | −0·061 |
| $r_3^*$ | 5·860 | 1·248 | 0·974 | 0·954 | 0·933 | −4·596* | −0·048 |
| $r_3-r_1$ | 0·091 | 0·138 | 0·842 | 0·643 | 0·463 | −8·497* | −0·296 |
| $r_3^*-r_1^*$ | 0·079 | 0·236 | 0·777 | 0·613 | 0·449 | −7·103* | −0·275 |
| $r_1-r_1^*$ | −0·158 | 0·275 | 0·550 | 0·518 | 0·438 | −7·391* | −0·447 |
| $r_3-r_3^*$ | −0·147 | 0·266 | 0·772 | 0·724 | 0·652 | −5·212* | −0·203 |

*Notes*:
Mean denotes the sample mean of the variables listed in column 1 in each panel, where $r_i(r_i^*)$ is the $i$th month Gensaki (Euroyen deposit) rates. Std denotes the sample standard deviation and $\rho(k)$ denotes autocorrelations of order $k$, for $k = 1, 2, 3$. DF(4) is the Dickey–Fuller test for a unit root in the following regression with an autoregressive process of order four:

$$\Delta x_t = c + \sum_{i=1,4} b_i \Delta x_{t-i} + a x_{t-1} + \varepsilon_t \qquad \text{for subperiods}$$

or

$$\Delta x_t = c_1 I\{t \leq \tau\} + c_1 I\{t > \tau\} + \sum_{i=1,4} b_i \Delta x_{t-i} + a x_{t-1} + \varepsilon_t \qquad \text{for a whole period}$$

where $I\{a\}$ is an indicator function whose value is equal to one if expression $a$ is true. The asterisk (*) denotes the rejection of the null hypothesis of a unit root (i.e. $a = 0$) at a 5% level. The 5% critical value is −2·86.

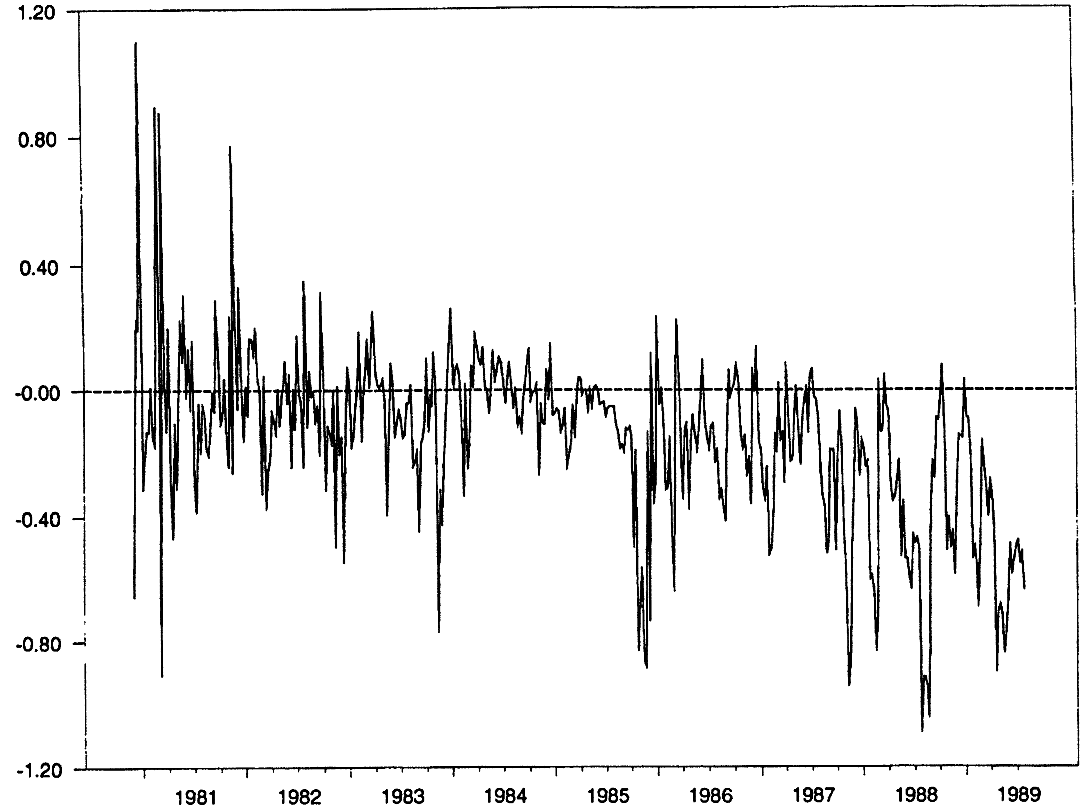


Figure 1. One-month Gensaki–Euroyen yield spread from December 1980 to September 1989

null hypothesis of a unit root cannot be rejected for four interest rates at least at a 5% significance level, but can be rejected for yield spreads at least at a 5% significance level (except that this hypothesis is rejected at a level slightly higher than 5% for the three-month Gensaki–Euroyen yield spread). The autocorrelation coefficients for the first three orders are reported. The results also reveal that the autocorrelation coefficients for the four interest rates die out very slowly, confirming that interest rates are consistent with the unit root hypothesis in the second period. Perron (1989) argues that the unit root hypothesis against trend stationary alternatives cannot be rejected by using standard DF tests if the true data-generating process is stationary with a one-time break in the trend. Hence, I include intervention dummies in the univariate autoregressive process in performing the DF tests. The null hypothesis of the unit root is still rejected, but autocorrelations for the four interest rates die out very slowly.[9]

### 3.2. The Number of Common Trends in Gensaki and Euroyen Deposit Rates

Examining the stationarity of the data and the number of common trends is necessary to provide evidence on the absence of apparent long-term profit opportunities for arbitrage across the two

[9] Other specifications of regressions used in the DF tests are also tried, including a trend in regression for subperiods, a break in a constant and a trend for the whole period, and different lags of the first differences of the data. The results are not affected.

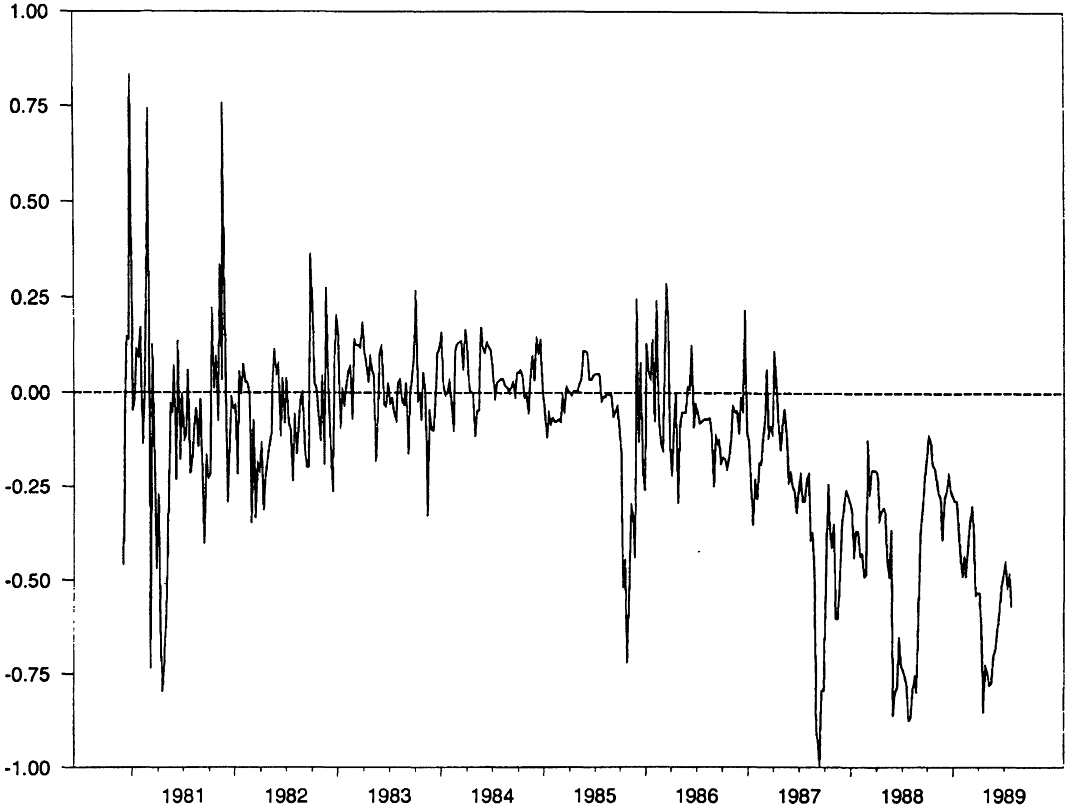


Figure 2. Three-month Gensaki–Euroyen yield spread from December 1980 to September 1989

markets. Kremers *et al.* (1992) suggest that with cointegration, tests based on the ECM statistics are more powerful than those based on the Dickey–Fuller test applied to the residuals of a static cointegrating relationship, because the latter test ignores valuable information about other variables in the system. Since the Johansen tests are likelihood ratio tests, a complete characterization of the data-generating process is unavoidable. The Schwarz (1978) criterion is employed to determine the lag length. Two lags in the VAR of the levels were chosen. I assume that the data in levels do not have a linear deterministic trend and a constant vector appears in cointegrating vectors.[10] Under this assumption, Table II presents statistics for the trace tests and the maximal eigenvalue tests for two subperiods and the whole period under unstable short-run dynamics, respectively. The hypotheses are denoted as $H_{q,u}$ and $H_{q,s}$ as shown in equations (7) and (8).

By using asymptotic critical values tabulated in Table A3 of Johansen and Juselius (1990), results in panels A and B in Table II reveal evidence on the rank of the cointegration matrix, $\Pi$, similar to that in Table I. The statistics for the trace tests show a rejection of the null hypothesis that the rank of the cointegration matrix is, at most, equal to zero, one, two, and three in the first period. The maximal eigenvalue test also rejects the null hypothesis that the rank of the cointegration matrix is equal to zero, one, two, and three. These results suggest that the rank of the cointegration matrix is equal to four and all four interest rate series are

[10] This hypothesis will be tested later.

Table II. Johansen test for the number of common trends under unstable short-run dynamics

| | Period 1 | | Period 2 | | Whole period | |
|---|---|---|---|---|---|---|
| $H_a$ | Trace | Eigen. | Trace | Eigen. | Trace | Eigen. |
| $q \leq 3$ | $19{\cdot}891^*$ | $19{\cdot}891^*$ | 4·270 | 4·270 | $25{\cdot}445^*$ | $25{\cdot}445^*$ |
| $q \leq 2$ | $73{\cdot}266^*$ | $53{\cdot}374^*$ | $23{\cdot}388^*$ | $19{\cdot}117^*$ | $83{\cdot}916^*$ | $58{\cdot}470^*$ |
| $q \leq 1$ | $136{\cdot}981^*$ | $63{\cdot}716^*$ | $77{\cdot}260^*$ | $53{\cdot}872^*$ | $166{\cdot}782^*$ | $82{\cdot}867^*$ |
| $q = 0$ | $231{\cdot}427^*$ | $94{\cdot}446^*$ | $142{\cdot}044^*$ | $64{\cdot}784^*$ | $268{\cdot}708^*$ | $101{\cdot}926^*$ |
| Autocorrelation of residuals: modified Ljung–Box LB(13) ($P$-values) | | | | | | |
| $r_1$ | 19·216 | (0·117) | 14·310 | (0·352) | 15·362 | (0·285) |
| $r_3$ | 13·698 | (0·395) | 10·193 | (0·678) | 14·080 | (0·368) |
| $r_1^*$ | 18·558 | (0·137) | 22·158 | (0·053) | 29·007 | (0·007) |
| $r_3^*$ | 12·571 | (0·506) | 13·455 | (0·413) | 13·690 | (0·396) |
| Hypothesis testing | | | | | | |
| Hypo | $H_{s.u}$ | | $H_u$ | | $H_\beta$ | |
| LR | 100·492 | (0·000) | 1·063 | (0·303) | 3·874 | (0·275) |

*Notes*:
$H_q$ denotes the null hypothesis regarding the rank of the cointegrating matrix, $q$, under the unstable short-run dynamics. Trace denotes the LR test for the null hypothesis that $q$ is at most, equal to $k$ (i.e. $q \leq k$) against the alternative hypothesis that $q = 4$ (the number of variables), while Eigen. denotes the LR test for the null hypothesis that $q$ is equal to $k$ against the alternative hypothesis that $q$ is equal to $k+1$, where $k = 0,\ldots,3$. As shown in Table A3 of Johansen and Juselius (1990), the 5% critical values for trace tests are 9·094, 20·168, 35·068, and 53·347 for $q = 3, 2, 1$, and 0, respectively. The 5% critical values for maximal eigenvalues are 9·094, 15·752, 21·894, and 28·167, respectively. The asterisk ($^*$) denotes the rejection of the null hypothesis at a 5% significance level.
MLB(13) denotes the modified Ljung–Box statistic which accounts for the heteroskedasticity of the data. $^*$ denotes the rejection of the null hypothesis at a 5% significance level.
$H_{s,u}$ denotes the null hypothesis for the stability of long-run parameters under the unstable short-run dynamics. The LR statistics for Hsu are distributed as a chi-square distribution with sixteen degrees of freedom. $H_\mu$ denotes the null hypothesis that the interest rates in levels do not have a linear deterministic trend in levels under the unstable short-run dynamics for the second period with the assumption that the data in levels for the first period have no linear trends. The LR statistic for $H_\mu$ is distributed as a chi-square distribution with one degree of freedom. $H_\beta$ denotes the null hypothesis for constraints in the cointegrating matrix $\beta = H\varphi$, where $H$ is set to be

$$H = \begin{bmatrix} 1{\cdot}0 & 1{\cdot}0 & 1{\cdot}0 \\ -1{\cdot}0 & 0{\cdot}0 & 1{\cdot}0 \\ 0{\cdot}0 & -1{\cdot}0 & 0{\cdot}0 \\ 0{\cdot}0 & 0{\cdot}0 & -1{\cdot}0 \end{bmatrix}$$

The LR statistic for $H_{\beta,u}$ is distributed as a chi-square distribution with three degrees of freedom.

stationary. In the second period, the null hypothesis that the rank of the cointegration matrix is no greater than (or equal to) zero, one, and two is rejected at a 5% level using the trace or maximal eigenvalue tests, while the null hypothesis that the rank of the cointegration matrix is no greater than (or equal to) four is not rejected at a 5% level. These results suggest that the rank of the cointegration matrix is no greater than three and hence, the off- and the onshore market interest rates have one common trend after August 1985.[11] In the whole sample period,

[11] Stock and Watson (1988) show that under the assumption of $k$ common trends, there are $q$ cointegration vectors, where $q = n - k$.

the null hypothesis that the rank of the cointegration matrix is at most equal to three is rejected.[12]

Since the Johansen test is a LR test, it depends on the fully specified autoregressive process for levels of data series. From Monte Carlo experiments, Cheung and Lai (1992) found that the Johansen test (1991) tends to find cointegration more often in the finite sample than in the asymptotic distribution and is more sensitive to the misspecification of lag length than to the non-normality of the errors. In light of this result, a careful selection of the lag length and the use of diagnostic testing for residuals are necessary to avoid overacceptance of cointegration. The middle part of Table I also reports the modified Ljung–Box test statistics of order 13 (MBL(13)) for the residuals from the VAR of the four interest rates, which investigates the null of no serial correlation. This statistic is derived using the generalized method of moments to account for heteroskedasticity (see West and Cho, 1993). Insignificant statistics of MBL(13) at a 5% level reveal no serial correlation in the regression residuals (except for one case in the one-month Euroyen deposit rate), suggesting that the lag length chosen for the Johansen test is appropriate.

The above results for the persistence of interest rates are in accord with those using the univariate Dickey–Fuller test in Table I as well as those in Campbell and Hamao (1992). The monetary policy in Japan has consistently utilized the interbank money rate as an operating variable and used window guidance to target bank credit. (See also Suzuki, 1990.) The non-stationary property of interest rates may be attributed to an easy monetary policy by decreasing an official discount rate after 1985 in Japan to prevent an adverse economic course from appreciation of the yen. Figure 3 plots fluctuations of the Gensaki rates and the official discount rate.

Three specification tests are also shown in the bottom part of Table II. The first LR test examines the stability of long-run parameters under unstable short-run parameters, the null hypothesis denoted as $H_{s,u}$. As shown in Quinto (1993), if the rank of the cointegration matrix for each subperiod is no less than that for the whole period, the LR test will have a chi-square distribution with sixteen degrees of freedom. LR test statistics in both panels show a rejection of stability of the long-run parameters and the rank of the cointegration matrix. While these results are interesting in regard to the persistence of interest rates, studies in the literature (e.g. Campbell and Hamao, 1992; Bonser-Neal and Roley, 1993) have not considered such a stability test. The second test examines the null hypothesis of whether the constant vector appears in the cointegration matrix and data in levels have a linear deterministic trend denoted as $H_{\mu}$. The statistics show no rejection for the second period. Because the Johansen test statistics in the first period reject that the rank of the cointegration matrix is, at most, equal to three, this hypothesis testing for the first period was not performed. The third hypothesis is that the yield spreads $s_t$ are the cointegrating vectors. This hypothesis puts linear restrictions on the cointegration matrix $\beta$, that is, $\beta = H\varphi$, where $\varphi$ was estimated under this alternative hypothesis. The matrix $H$ is

$$H = \begin{bmatrix} 1{\cdot}0 & 1{\cdot}0 & 0{\cdot}0 \\ -1{\cdot}0 & 0{\cdot}0 & 1{\cdot}0 \\ 0{\cdot}0 & -1{\cdot}0 & 0{\cdot}0 \\ 0{\cdot}0 & 0{\cdot}0 & -1{\cdot}0 \end{bmatrix}$$

[12] I also performed the Johansen test under stable short-run parameters with intervention dummies to examine the robustness of the results for the rank of the cointegration matrix. In this case, the size of the Johansen test is affected in the presence of intervention dummies. By using the DISCO program provided by Johansen and Nielsen (1993) with 10,000 replications to obtain the simulated critical values, I found that the results are consistent with those under unstable short-run parameters.

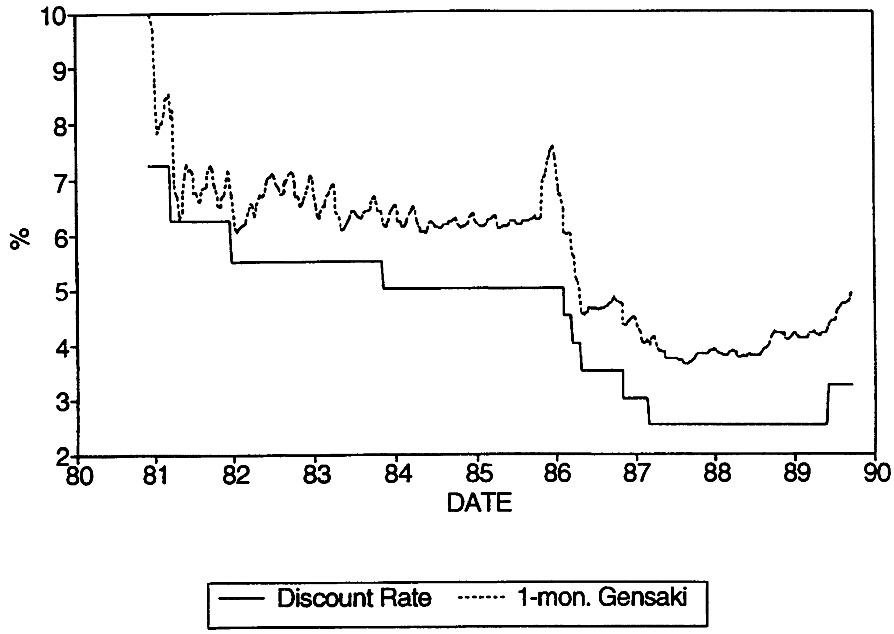


Figure 3. One-month Gensaki rate and discount rate from December 1980 to September 1989

The test statistic is distributed as a chi-square distribution with three degrees of freedom. The test statistic for the second period shows no rejection of the null hypothesis at a 5% significance level. As a whole, I can draw the following conclusions from this analysis. First, Euroyen deposit and Gensaki rates were stationary before the deregulation and there was one common trend in Euroyen deposit and Gensaki rates, indicating no long-run apparent profit opportunity after deregulation. Second, there is a structural break in the rank of the cointegration matrix or the long-run parameters.

## 4. AN ERROR CORRECTION MODEL

This section examines the predictive power of cross-market information on interest rates and the dynamic path of the yield spreads in response to a shock. To address these two issues, the following two procedures are employed. First, an error-correction model is estimated and tests for no cross-market causality, for the unpredictability of Gensaki–Euroyen yield spreads, and for no structural break in the short-run parameters of interest rate processes are reported in Section 4.1. Second, the impulse response functions for Gensaki–Euroyen yield spreads to one-unit shock in one-month Gensaki and Euroyen deposit interest rates, as well as their confidence intervals for the periods before and after August 1985, are plotted in Figures 4 and 5.

### 4.1. Estimated Results for An Error-correction Model

Although the rank of the cointegrating matrix is different in the first and the second periods, a root of the characteristic function for the VAR is very close to a unit circle. In order to compare results for cross-market causality and impulse response functions in both periods, I use the same ECM specification (i.e. equation (6)) with constraints on the cointegration matrix. Most studies

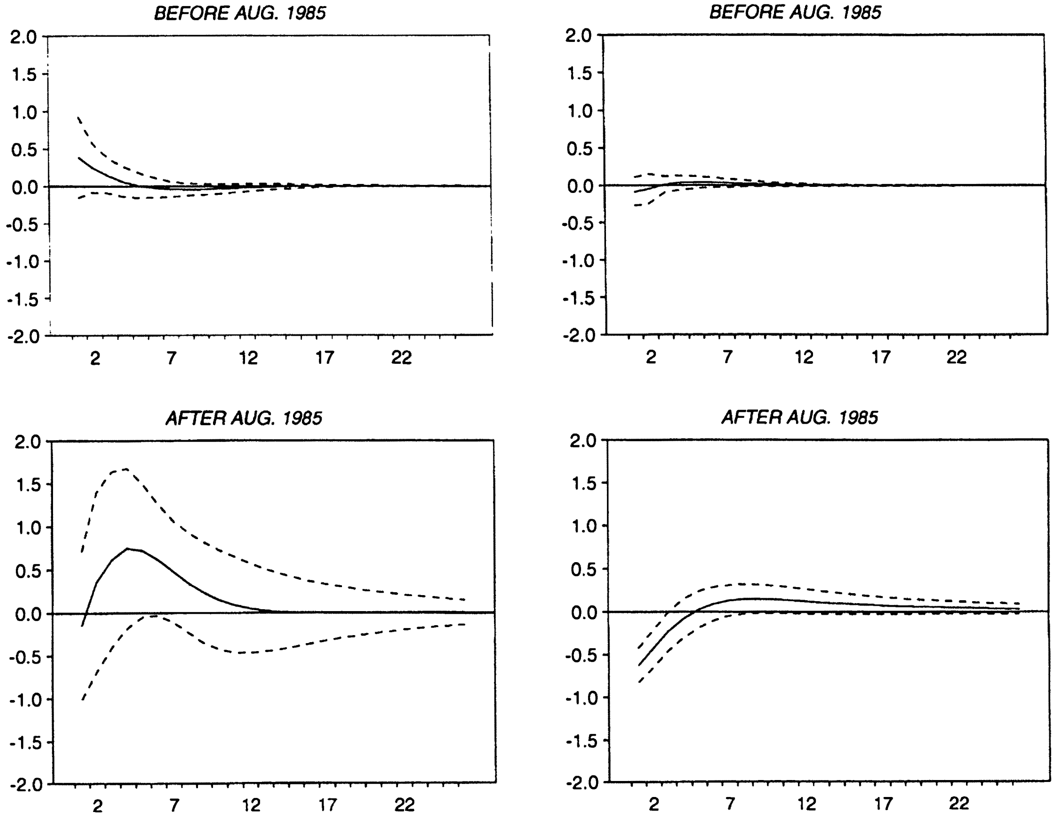


Figure 4. The impulse response curve for the one-month yield spread to a one-unit shock in the one-month Gensaki and Euroyen rates

attempt to provide estimation methods to improve the finite sample biases in the estimators of cointegrating vectors (see Phillips and Loretan, 1991, for finite sample properties of alternative estimators). Because cointegrating vectors are constrained at the theoretical values as in equations (2) to (4), there is no need to estimate the cointegration vector. Moreover, the hypothesis testing $H_\mu$ in Section 3.2 suggests that the constant vector in the ECM lies in the space spanned by the adjustment matrix. This gives cross-equation constraints on the adjustment matrix and the constant vector. Due to the above two reasons, a multivariate nonlinear OLS with these cross-equation constraints is used. The estimated coefficients for the error-correction model with heteroskedasticity-consistent standard errors (see White, 1984) are presented for the first and the second periods in Table III.

Some conclusions emerge from Table III. First, $R^2$ shows that changes in lagged interest rates and yield spreads can explain between 20% and 55% of the total variations of changes in the four interest rates in the first period and between 9% and 28% of the variations in the second period. Second, there is some weak evidence that the offshore market information such as changes in Euro yen deposit rates (denoted as Wald($\Delta r^*$)) affects changes in Gensaki rates, whereas the onshore market information (denoted Wald($\Delta r$)) also affects changes in Euroyen deposit rates. Third, the Gensaki yield spread between one and three months helps to predict the one-month Gensaki rate for both periods, which is one of the implications in the rational

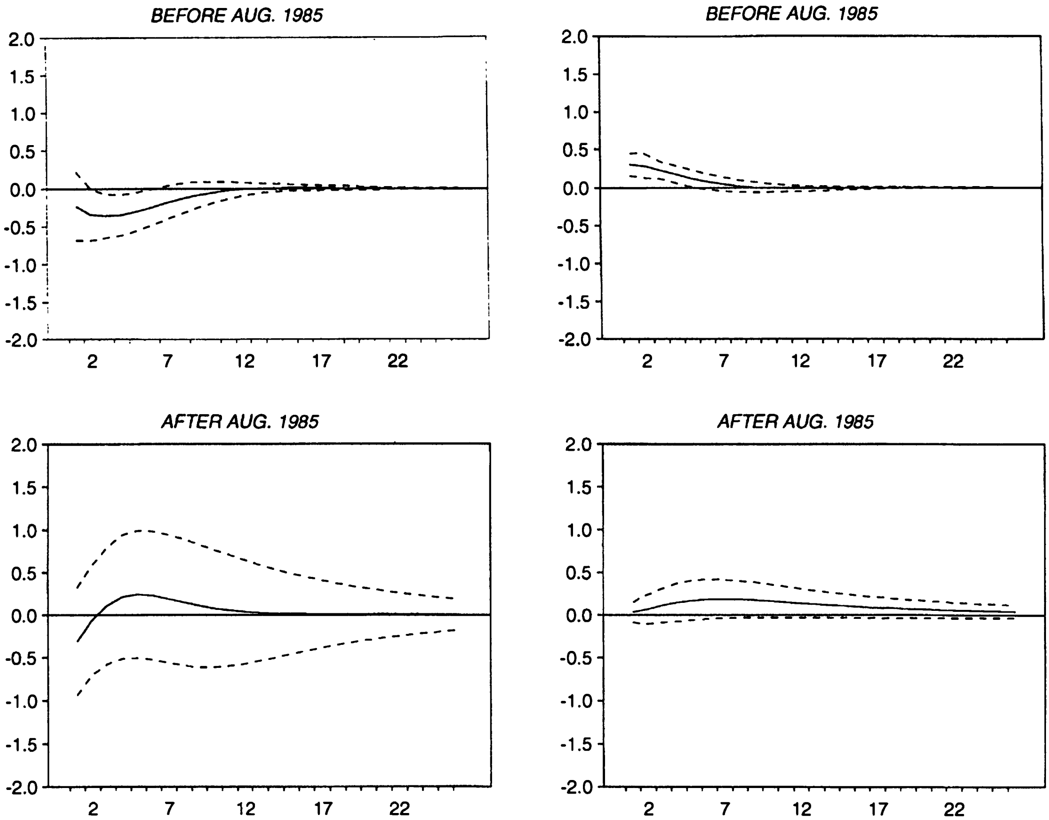


Figure 5. The impulse response curve for the three-month Gensaki–Euroyen yield spread to a one-unit shock in the one-month Gensaki and Euroyen rates

expectations hypothesis. In contrast to Campbell and Hamao (1992), the Gensaki yield spread between one and three months significantly explains the variations in the one-month Gensaki rate before August 1985. This finding is analogous to the evidence presented for US Treasury bills and bond markets by Fama (1984) and Campbell and Shiller (1987), who found that yield spreads help predict future short-term interest rates.

To examine the second and third conditions of market linkage, tests for no Granger-causality, for the unpredictability of Gensaki–Euroyen yield spreads, and for the stability of the short-run parameters and Gensaki–Euroyen yield spreads are reported in Table IV. Wald test statistics for no cross-market causality are reported in panel A. These statistics indicate that the null hypotheses, $H_G$, are rejected, suggesting a bi-directional Granger causality relation between the Gensaki and Euroyen deposit rates. Panel B reports Wald tests for unpredictability of one- and three-month Gensaki–Euroyen yield spreads. These suggest that Gensaki–Euroyen yield spreads are predictable except for the one-month Gensaki–Euroyen yield spread in the first period. Panel C presents tests for no structural break in the short-run parameters $\Gamma$ (in the first four columns) for individual interest rate processes and for no structural break in $\delta$ (in the last column). The Wald tests statistics reject these null hypotheses. Comparison of the estimates of $\delta$ before and after deregulation shows that the Gensaki–Euroyen yield spreads decreased after deregulation.

Table III. Estimation of error correction models

| | $\Delta r_1$ Coeff. (Tstat) | $\Delta r_3$ Coeff. (Tstat) | $\Delta r_1^*$ Coeff. (Tstat) | $\Delta r_3^*$ Coeff. (Tstat) | $\delta$ Coeff. (Tstat) |
|---|---|---|---|---|---|
| A. Before deregulation | | | | | |
| $s_{1,t-1}$ | 0·354 | 0·139 | 0·461 | 0·148 | 0·143 |
| | (3·962) | (3·425) | (3·559) | (1·774) | (7·745) |
| $s_{2,t-1}$ | −0·199 | −0·086 | 0·669 | 0·290 | −0·048 |
| | (−2·594) | (−1·449) | (3·340) | (2·332) | (−3·899) |
| $s_{3,t-1}$ | 0·141 | −0·072 | 0·191 | 0·193 | −0·027 |
| | (1·239) | (−1·202) | (1·051) | (1·667) | (−1·735) |
| Wald($\Delta r$) | 9·448 | 14·037 | 1·294 | 3·799 | |
| | (0·051) | (0·007) | (0·862) | (0·434) | |
| Wald($\Delta r^*$) | 11·377 | 13·634 | 30·028 | 13·288 | |
| | (0·023) | (0·009) | (0·000) | (0·010) | |
| $R^2$ | 0·523 | 0·553 | 0·324 | 0·203 | |
| B. After deregulation | | | | | |
| $s_{1,t-1}$ | 0·771 | 0·491 | 1·312 | 0·750 | 0·036 |
| | (4·134) | (3·403) | (2·691) | (2·820) | (5·110) |
| $s_{2,t-1}$ | −0·100 | −0·049 | 0·304 | 0·033 | −0·336 |
| | (−2·818) | (−1·583) | (2·468) | (0·369) | (−7·660) |
| $s_{3,t-1}$ | 0·025 | 0·031 | −0·193 | 0·010 | −0·348 |
| | (0·812) | (−1·086) | (−1·767) | (0·119) | (−6·847) |
| Wald($\Delta r$) | 4·111 | 4·563 | 3·267 | 9·032 | |
| | (0·391) | (0·335) | (0·514) | (0·060) | |
| Wald($\Delta r^*$) | 4·716 | 14·155 | 20·787 | 0·883 | |
| | (0·318) | (0·007) | (0·000) | (0·927) | |
| $R^2$ | 0·279 | 0·226 | 0·233 | 0·093 | |

*Notes*:
A least square estimation with cross-equation constraints on $\delta$ and $\alpha$ is applied to the following ECM:

$$\Delta y_{i,t} = \sum_{j=1,4} \Gamma_{ij}\Delta y_{j,t-1} + \sum_{j=1,3} \alpha_{ij}(s_{j,t-1} - \delta_j) + u_t \qquad \text{for } i = 1, \ldots, 4$$

The second to fourth columns report the estimate of $\alpha_{ij}$, the coefficient of $s_{jt-1}$, (as listed in the first column) on the regression of $y_{i,t}$, (as listed in the first row), for $j = 1,\ldots,3$ and $i = 1,\ldots,4$. The last column reports the estimate of $\delta_j$ for $s_{j,t-1}$.
Wald($\Delta r$) denotes a Wald test for the null hypothesis that $\Gamma_{i1} = \Gamma_{i2} = 0$ and Wald($\Delta r^*$) denotes a Wald test for the null hypothesis that $\Gamma_{i3} = \Gamma_{i4} = 1\ 0$. Both are distributed as a chi-square distribution with two degrees of freedom.

The above analyses suggest that the interest rates in these two markets are interdependent, but on- and offshore yield spreads enlarged and became more predictable after deregulation. Such predictability may be attributed to the existence of differences in either risk or transaction cost.

## 4.2. Impulse Response Function

To further examine how fast the Gensaki–Euroyen yield spreads can adjust to new information in one market, Figures 4 and 5 plot the impulse response functions of one- and three-month Gensaki–Euroyen yield spreads to a one-unit shock in one-month Gensaki (right) and Euroyen deposit rates (left), and their 5% confidence intervals. The one-month Gensaki–Euroyen yield spreads significantly respond only to a one-unit shock in Euroyen deposit rates after August 1985. The responses of the three-month Gensaki–Euroyen yield spread to a one-unit shock in

Table IV. Hypothesis tests

A. Tests for no cross-market causality

| | $r_1$ | $r_3$ | $r_1^*$ | $r_3^*$ |
|---|---|---|---|---|
| Period 1 | | | | |
| $H_{G,r}$ | | | 69·695 | 37·154 |
| | | | (0·000) | (0·000) |
| $H_{G,r^*}$ | 30·826 | 44·707 | | |
| | (0·000) | (0·000) | | |
| Period 2 | | | | |
| $H_{G,r}$ | | | 30·757 | 11·196 |
| | | | (0·000) | (0·000) |
| $H_{G,r^*}$ | 14·276 | 16·194 | | |
| | (0·006) | (0·003) | | |

B. Tests for the unpredictability of Gensaki–Euroyen yield spreads

| | Wald($s_2$) | ($P$-value) | Wald($s_3$) | ($P$-value) |
|---|---|---|---|---|
| Period 1 | 6·177 | (0·519) | 71·147 | (0·000) |
| Period 2 | 292·240 | (0·000) | 801·106 | (0·000) |

C. Tests for no structural break

| $r_1$ | $r_3$ | $r_1^*$ | $r_3^*$ | $\delta$ |
|---|---|---|---|---|
| 16·442 | 17·520 | 22·307 | 9·849 | 42·985 |
| (0·002) | (0·002) | (0·000) | (0·043) | (0·046) |

*Notes*:
A nonlinear least square estimation is applied to the following ECM:

$$\Delta y_{i,t} = \sum_{j=1,4} \Gamma_{ij}\Delta y_{j,t-1} + \sum_{j=1,3} \alpha_{ij}(s_{j,t-1} - \delta_j) + u_t \qquad \text{for } i = 1, \ldots, 4$$

Panel A reports test statistics for no Granger-causality. The null hypotheses—$H_{G,r}$: $\Gamma_{1j} = \Gamma_{2j} = \alpha_{1j} = \alpha_{2j} = \alpha_{3j} = 0$ for $j = 3, 4$ and $H_{G,r^*}$: $\Gamma_{3j} = \Gamma_{4j} = \alpha_{2j} = \alpha_{3j} = 0$ for $j = 1, 2$—are listed in the first column, and are distributed as a chi-square with five and four degrees of freedom, respectively. Panel B reports the unpredictability test for the null hypothesis that $h'\Gamma = 0$ and $h'\alpha = -1$, where $h' = [1,0,-1,0]$; it is distributed as a chi-square distribution with seven degrees of freedom. Panel C reports Wald tests for no structural break, which is a chi-square distribution with seven (three) degrees of freedom for short-run parameters $\Gamma$ (the constant vector in the cointegration matrix, i.e. $\delta$).

Euroyen and Gensaki interest rates are significant before deregulation, but not afterwards. Although the responses of one- and three-month Gensaki–Euroyen yield spreads to a one-unit shock in one-month Euroyen and in Gensaki rates seem to increase after deregulation, larger standard errors lead to an insignificant response in most cases.[13]

[13] The test for the unpredictability of Gensaki–Euroyen yield spreads is a joint test for the hypothesis that all coefficients in the regression of Gensaki–Euroyen yield spreads are statistically equal to zero. The test for no response of Gensaki–Euroyen yield spreads is a test for the hypothesis that the nonlinear combinations of those coefficients are jointly equal to zero.

Table V. Wald tests for no effect on the impulse response function

| | Response to $r_1$ | | Response to $r_1^*$ | |
|---|---|---|---|---|
| $(m^*,m^*)$ | $s_2$ | $s_3$ | $s_2$ | $s_3$ |
| A. Before deregulation | | | | |
| 1,4 | 3·802 | 7·972 | 4·533 | 23·678 |
| | (0·433) | (0·093) | (0·339) | (0·000) |
| 5,8 | 3·455 | 9·482 | 2·467 | 23·167 |
| | (0·485) | (0·050) | (0·651) | (0·000) |
| 9,12 | 0·068 | 0·971 | 4·822 | 14·878 |
| | (0·999) | (0·041) | (0·306) | (0·005) |
| B. After deregulation | | | | |
| 1,4 | 4·918 | 2·672 | 70·133 | 7·063 |
| | (0·296) | (0·614) | (0·000) | (0·133) |
| 5,8 | 4·081 | 3·679 | 34·513 | 4·006 |
| | (0·395) | (0·451) | (0·000) | (0·405) |
| 9,12 | 5·353 | 2·734 | 5·891 | 3·443 |
| | (0·253) | (0·603) | (0·207) | (0·487) |

*Notes*:
Wald test statistics for no response of $s_{t+8}$ to $r_{1,t}(r^*_{1,t})$ for all $s = m^*, m^* + 1, \ldots, m^*$ are reported and distributed as a chi-square with four degrees of freedom.

Table V reports the Wald tests for the null hypothesis that the response of the Gensaki–Euroyen yield spread for lags $m_.$ to $m^*$ are jointly equal to zero. The joint tests suggest that the response of the one-month Gensaki–Euroyen yield spread to a one-unit shock in the one-month Euroyen deposit rate dies out eight weeks later after deregulation but dies out very quickly before deregulation. The response of the three-month Gensaki–Euroyen yield spread to a one-unit shock in the one-month Gensaki and Euroyen deposit rates dies out more slowly before deregulation than afterwards.

## 5. CONCLUSIONS

This paper examined whether Japan's financial deregulation decreased the depth of the Gensaki market and weakened the linkage of the Gensaki and Euroyen deposit markets. I measured the linkage of these two markets as the adjustment speed of Gensaki–Euroyen yield spreads to a shock and the usefulness of cross-market information in predicting interest rates. To examine their linkage, I employed the Johansen test, estimated the ECM model, and constructed the impulse response function. The results suggest that there are no apparent long-term profits for arbitrage between the two markets; that information transmits across the two markets; and that the degree of linkage between the two markets becomes weaker.

A weaker linkage between Gensaki and Euroyen deposit rates may be attributable to the relative lower liquidity in Gensaki trading as compared to other markets Such a decline may be due to the emergence of new assets, deregulations of Japan's CDs market, and the existence of securities transfer taxes in the Gensaki market. This paper, in accordance with the implications of Bonser-Neal and Roley (1993), suggests that the Gensaki market is not a representative money market in Japan. Since regulations on the issuance, size, and maturity of CDs have been gradually removed, the CD (with its repurchase agreement market) has suppressed the Gensaki and become the most active money market. The linkage of the Japanese CD, Euroyen CD, and US CD will be left for future research.

## ACKNOWLEDGEMENTS

I would like to thank Mark Watson (the editor), two anonymous referees, Michael T. K Horvath (a referee), Carmela Quinto, and participants in the 1993 Econometric Society summer meetings for helpful comments. I am also grateful to Ken West, Don Hester, Takatoshi Ito, Heather Anderson, Steve Durlauf, Peter Knez, and Mico Loretan for helpful discussions; and to Takatoshi Ito, Yoshiaki Shikano, Ken West, and the Japan Securities Dealers Association for providing me with the data used in this paper. Financial support from the Graduate School of the University of Wisconsin-Madison and computational assistance from Hsiao-Hua Leng and Ruey Yau are also gratefully acknowledged. All remaining errors are mine.

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