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Interpreting Changes in the Volatility of Yields on Japanese Long-term Bonds

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This paper investigates empirically the relations between changes in volatilities of holding period returns on JGBs and changes in U.S. interest rates and the yen/dollar exchange rate. Weekly and quarterly holding period returns are constructed for the period March 1986 through May 1988. Then quadratic and Fourier series approximations to the conditional variances of these yields are estimated. Significant variation in the conditional variances of holding period returns is documented. These results are interpreted in the light of the changing patterns of trading volumes and capital flows between the U.S. and Japan in the 1980s.

I. Introduction

The liberalization of financial markets in Japan during the past decade has contributed to a much broader participation by domestic and overseas investors in Japanese long-term government bond (JGB) markets. The associated increases in trading volumes and international capital flows have, in turn, increased the influence of global economic developments on the time-series properties of yields in these markets. In this paper, I investigate empirically the relative contributions of various domestic and foreign financial variables to the temporal changes in the levels and volatilities of Japanese long-term bond yields. Particular attention is given to the correlations between changes in the volatility of bond yields in Japan and shifts in the term structure of U.S. Treasury bond yields and the yen/dollar exchange rate.

The linkages between interest rates in Japanese, European, and U.S. bond markets have been studied recently by the Bank of Japan (1986, 1988), Bomhoff and Schotman (1987), and Kool and Tatom (1988). Bank of Japan (1986, 1988) and Kool and Tatom

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(1988) estimated best linear predictors of bond yields based on contemporaneous and lagged values of Japanese interest rates and other domestic and foreign macroeconomic variables. Bomhoff and Schotman (1987) study a version of an ARCH-M model (Engle, Lilian and Robbins 1987) of Japanese bond yields.

The analysis of Japanese bond markets in this paper differs in several important respects from these previous studies. First, this study focuses primarily on changes in volatility instead of the conditional mean of returns. Volatility in holding period returns is measured as the variance of the holding period return conditional on a set of economic variables known at the date the investment is undertaken. Thus, the perspective is that of a forward looking investor using current and lagged information to forecast the future. Both domestic and international sources of conditional volatility are examined, including past changes in Japanese and U.S. short- and long-term rates and changes in the yen/dollar exchange rate.

The first step in calculating an estimate of volatility is the estimation of the best predictor of a holding period return conditioned on these variables. In order that a wide variety of possible representations of this best forecast can be accommodated, it is modeled using a Fourier series approximation (Gallant 1982; Pagan and Hong 1988) to a general functional form. This representation permits nonlinear conditional mean functions. In particular, it accommodates the ARCH-M like relations studied by Bomhoff and Shotman (1987), without restricting *a priori* the conditional mean of excess returns to be proportional to the conditional variance. Next, the prediction error implied by this best forecast is calculated. Finally, the conditional variance of the holding period return is modeled as the best forecast of the square of this forecast error, with forecasts being based on a Fourier functional form. The conditioning information is lagged squared forecast errors and the macroeconomic information used to forecast the levels of returns.

The measurement of returns and the sampling interval of the data is also notably different from previous studies. Previous investigations of long-term Japanese bond returns have typically focused on post-1978, monthly data on yields-to-maturity. In contrast, I study weekly data on holding period returns over a sample period beginning in 1986. Holding period returns are the *ex post* investment yields obtained in the bond markets and, therefore, the risk characteristics of these yields are more relevant for portfolio decisions than is the volatility of yields to maturity. Two different holding period returns are examined: the return on the bond with maturity closest to ten years and the return on the current "benchmark" ten year bond. The institutional considerations outlined subsequently suggest that the volatilities of these two return series may be quite different. The returns are calculated for both one week and three-month holding periods.

The sample period is restricted to the relatively short 1986–88 period in an attempt to control for some of the fundamental changes that occurred in Japanese financial markets over the past decade. There were several developments in 1984 and 1985 in particular that lead to significant changes in the trading patterns of Japanese investors in both domestic and foreign bond markets. I argue subsequently that these changes may explain the recent increases in the correlations between returns on JGBs and changes in foreign interest rates and exchange rates documented, for instance by the Bank of Japan (1988). In the variance decompositions associated with the monthly vector autoregressions fit by the Bank of Japan (1988) for the sample period January 1978 through March 1983, over half of the variation in the Japanese long-term yield-to-maturity was due to its own past history. In contrast, for the sample period April 1983 through June 1988, this percentage declined to 22 percent, with about 86 percent of the variation in JGB yields being attributable to variation in the U.S. long-term rate and the exchange rate. This paper explores the linkages between the volatility of holding period returns and foreign financial variables over the recent period of increased integration and capital flows.

The focus will be primarily on the yields on long-term bonds. One reason for this focus is that the long-term market is more developed, less regulated, and has a much larger trading volume by both domestic and foreign investors than the short-term bond markets in Japan (see Section II). In addition, the economic reasons for the sympathetic movements in short- and long-term rates that underlie the expectations-based theories of the term structure commonly applied to the U.S. government bond markets are much weaker (and in some cases virtually absent) in Japan. Thus, while changes in short-term yields may be correlated with the temporal behavior of yields in long-term markets, a preliminary exploration of the volatility of bond yields without the imposition of the structure of specific term structure models seems warranted.[1]

The remainder of this paper is organized as follows. In Section II changes in the regulations affecting trading in Japanese bond markets are related to recent trading patterns by Japanese investors in both the domestic and foreign bond markets. The high turnover ratios of domestic and foreign bond holdings documented in this section suggest that many investors have very short investment horizons. These patterns are used to motivate the choice of conditioning variables underlying the specification of the conditional variance of holding period returns on bonds. The econometric model of conditional means and variances is set forth in more detail in Section III. The empirical results are presented in Section IV, and concluding remarks are presented in Section V.

II. Historical Background

Though the liberalization of domestic bond markets in Japan has narrowed substantially the organizational differences between markets in Japan, the U.S., and Europe,

[1] Dunn and Singleton (1986) and Singleton (1988) derive term structure models relating the volatilities of short- and long-term yields under the assumption that marginal rates of substitution of consumption and real bond yields are jointly log-normally distributed. The validity of the assumptions about trading opportunities underlying these derivations seems questionable for Japan.

significant differences in legal restrictions on trading and in trading practices by large institutional investors across countries remain. In this section I discuss some implications of these differences, as well as several recent regulatory changes, for modeling changes in interest rates on Japanese government bonds.[2]

Prior to 1974, there was a small stock of outstanding government debt and a correspondingly low transactions volume in the secondary markets for government debt. Furthermore, only ten year bonds were issued and these bonds were priced by a syndicate in consultation with the Bank of Japan. Bonds were typically bought by an underwriting syndicate of financial institutions at prices above the prices of securities in the secondary market. There was an (implicit) understanding that the Bank of Japan would purchase these bonds after a period of one year and in practice such purchases were frequently consummated (Takagi 1988). Banks were not allowed to resell these bonds in the secondary market.[3]

The slowdown in economic growth after 1974 resulted in large issues of government bonds to finance the growing budget deficits. As a consequence of the substantial increase in government bonds held by banks, the Bank of Japan suspended its guarantee to repurchase these securities. This, in turn, contributed to a significant deterioration in the liquidity of the investment portfolios of the members of the syndicate. Consequently, in April 1977, financial institutions were allowed to sell government bonds in the secondary market after a minimum holding period of one year. The secondary sales of bonds by banks after holding securities for 100 days were authorized during 1981. And over-the-counter sales of newly issued bonds were permitted starting in October 1983. The last two changes also apply to the sale of intermediate term government bonds (maturities of 2, 3, and 4 years), which were issued for the first time between 1978 and 1980.

In June, 1984 bank dealing in bonds was authorized. Initially, 34 banks were allowed to trade as dealers in bonds, and dealing was restricted to bonds with less than two years remaining to maturity. Bank dealings in bonds were completely liberalized in June 1985. At this time banks were also allowed to use the interdealer brokerage services provided by the Nihon Sogo Shoken.[4] Throughout this period the composition of the bonds being traded was evolving and the number of days following the subscription of a bank to a bond issue that had to elapse before secondary trading could begin changed several times. Currently, secondary trading for the dealing account can begin immediately after

[2] Excellent descriptions of the evolution of financial markets during the past fifteen years can be found in Suzuki (1987), Bank of Japan (1986), and Takagi (1988).

[3] Secondary trading in government bonds on organized securities exchanges commenced in October 1966. There were no restrictions subsequently on trading by individuals and life insurance companies of the small amount of government bonds held by these agents (Kuroda 1982).

[4] The Nihon Sogo Shoken is a "broker's broker" that processes anonymously trades for dealers. All of the dealers trading with the Nihon Sogo Shoken (142 securities companies and 184 banks as of April 1989) can see a list of up-to-the-second quotes (bids and asks).

(starting on the day of) each issue; and for the investment account after ten days following the issue date.

With the admission of banks as dealers in 1985 came a substantial increase in volume in the over-the-counter markets for long-term government bonds; see the graph of over-the-counter (OTC) turnover of interest bearing JGBs displayed in Figure 1. Since the middle of 1985, the monthly turnover of JGBs in the OTC market has exceeded the outstanding balance of JGBs in all but four months. Accompanying the increased volume was a notable decline in transactions costs, as measured by the bid-ask spread, in the OTC markets for bonds (Takagi 1987). The spreads are now comparable or smaller than those on the corresponding government bonds in the U.S. (Bank of Japan 1986).

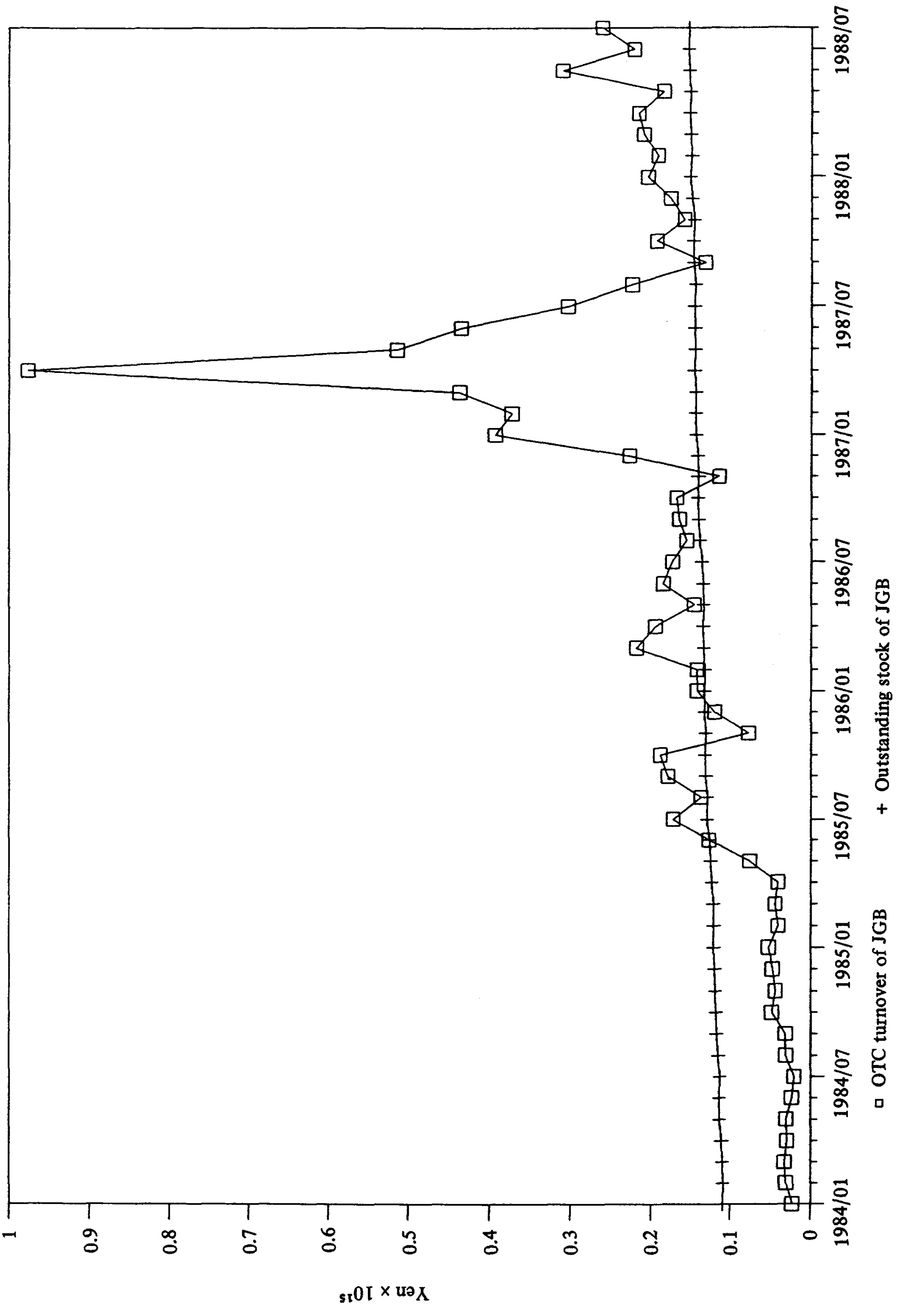
Most of the trading in the OTC market for government bonds has been between dealers. As recently as 1983, transactions by securities companies and banks through their dealing accounts represented only 30% of secondary market transactions. By 1985, this percentage had increased to 70% (Bank of Japan 1988). The total transactions volume of securities companies and banks in their portfolio and dealing accounts represents about 90% of all secondary market transactions. In other words, the secondary market is effectively an inter-dealer market for these financial institutions.

Furthermore, a striking feature of the Japanese bond market is that most of this turnover has been concentrated in a single "benchmark" issue. Figure 1 grossly understates the turnover ratio for the benchmark, since the benchmark issue has the largest trading volume among long-term bonds and at the same time is a small percentage of the outstanding stock of JGBs. Indeed, the percentage of the total trading volume through the broker's broker market accounted for by the benchmark issue 105 ranged between 90% and 95% in late 1988. Furthermore, the daily volume during October, 1988 was typically between ¥1 and ¥2 trillion and the outstanding stock of the benchmark was about ¥1. 5 trillion, so the stock often turned over more than once per trading day.

Because of the high "liquidity" in this market relative to the OTC markets for other intermediate and long-term government bonds, the benchmark issue trades at a substantial premium relative to non-benchmark issues. The spread has been 5 or 6 times larger than the typical spread in the U.S. between "on" and "off the run" government bonds (e.g., Sargen, Schoenholtz, Blitz and Elhabashi 1986). During the sample period for this study, March 1986–May 1988, three different bond issues served as the benchmark issue (see Section III for details). When a benchmark issue looses its status as the benchmark bond, then the volume of trade declines substantially as does the premium reflected in the price. Concurrently, the price of the new benchmark bond adjusts to reflect the increased liquidity associated with the increase in its trading volume. Typically, the transition period lasts for about a month, during which time price volatility appears to increase with the uncertainty regarding which of the candidate issues will become the next benchmark (Sargen, et. al. 1986).

There is an institutional feature of the Japanese financial markets that seems to

Figure 1. Turnover and Stock of JGB



reinforce the presence of this benchmark phenomenon. Specifically, the payment by insurance companies of dividends to policy holders from capital gains is limited under current law. And until recently only coupon income could be used for dividends. These regulations encourage a buy-and-hold strategy on the part of insurance companies for bonds with relatively high coupon yields. Thus, bonds that are being traded at a discount relative to otherwise comparable bonds are bought by insurance companies and some pension fund managers. This behavior may be contributing to the low volume in non-benchmark issues.

While institutional investors pursued a buy-and-hold investment strategy for portions of their fixed income portfolios, other portions of their portfolios experienced relatively high turnover ratios. In particular, starting in the early 1980s insurance companies had high turnover ratios on their total bond portfolios compared to many other institutional investors, as well as relatively large investments in foreign bonds (Bank of Japan 1982). The investment patterns by insurance companies in Japanese government bonds during the period 1984–88 are displayed in Figure 2. Notice that there has been an upward trend in both buy and sell transactions by insurance companies since mid-1985.

Interestingly, there were corresponding increases in the turnover ratios for foreign investors in Japan and Japanese investors in foreign countries. Regarding the former, the Bank of Japan (1986) estimated that from 1981 to 1984 the number of turnovers per year in the total government bond portfolios of overseas investors in Japan increased from 2 to 6. Overall, however, the dollar value of the investments in Japanese bonds by overseas investors was on average less than 25% of the value of the investments in foreign bonds by Japanese investors between 1986 and 1988.[5]

The buy and sell transactions in foreign bonds by Japanese investors, on monthly basis, are displayed in Figure 3. What is perhaps most striking about Figure 3, is the close relation between buy and sell transactions. The net acquisitions of bonds were small indicating that most of the trading was associated with rapid turnover of bonds. Furthermore most of the trading activity was concentrated in the U.S. markets. In 1985, about 85% of the turnover of foreign bond holdings by Japanese investors was trading activity in the U.S., while 51.4% of the total portfolio of bonds was comprised of U.S. securities. In contrast, in June 1988, about 35% of the stock of foreign bond holdings by Japanese was U.S. bonds, yet over 90% of the turnover of their portfolios was attributable to trading in the U.S. bond markets. Evidently, while the international bond portfolios of the Japanese have become more diversified, the concentration of trading activity has increased.

These changes in trading patterns both by Japanese investors in the U.S. and overseas investors in Japan occurred around the times of the development of the benchmark

[5] For the period 1986–88, the volume of investments by overseas investors in the Japanese market was typically less than $50 billion.

Figure 2. Sales and Purchases of JGB by Japanese Insurance Companies

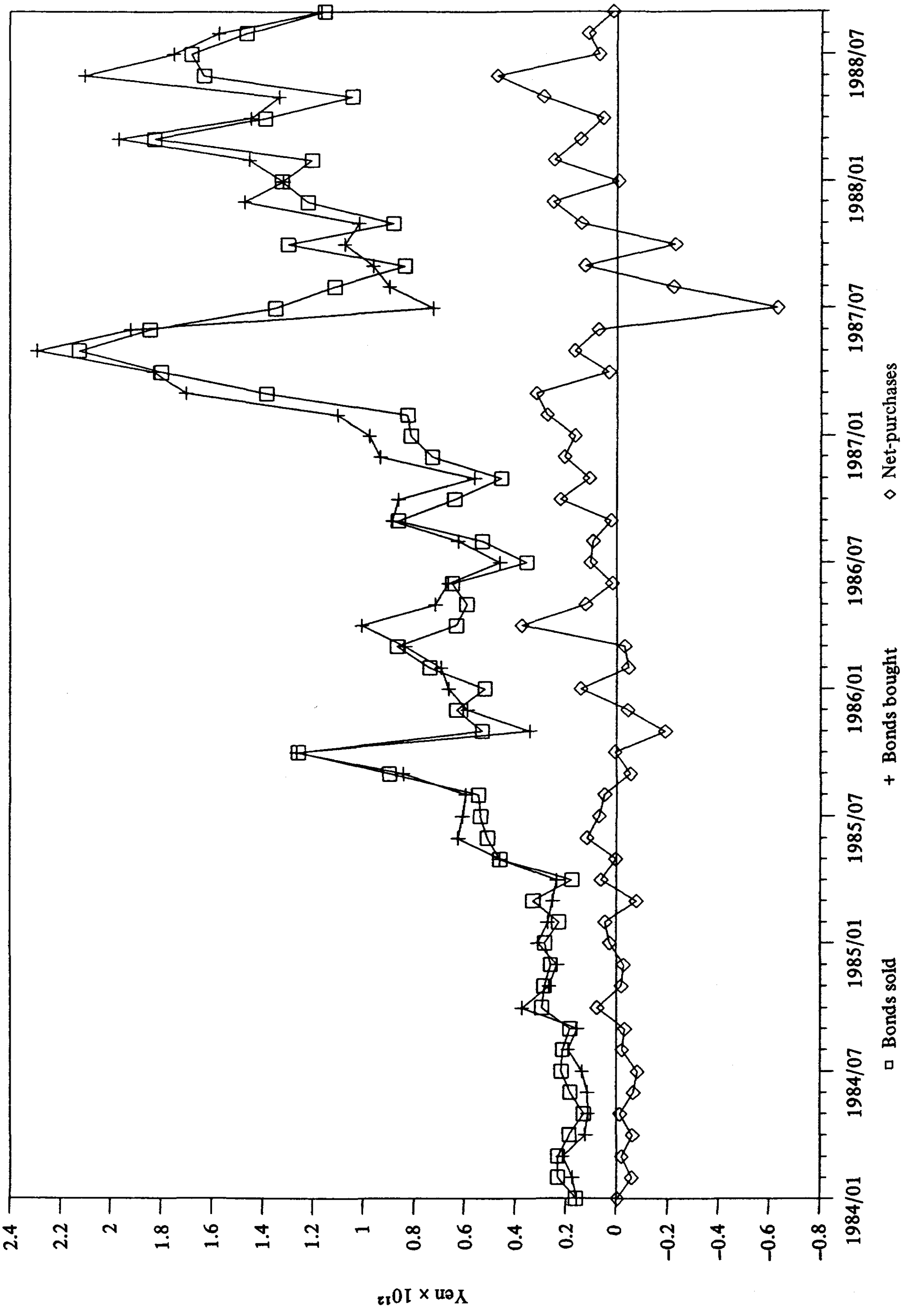
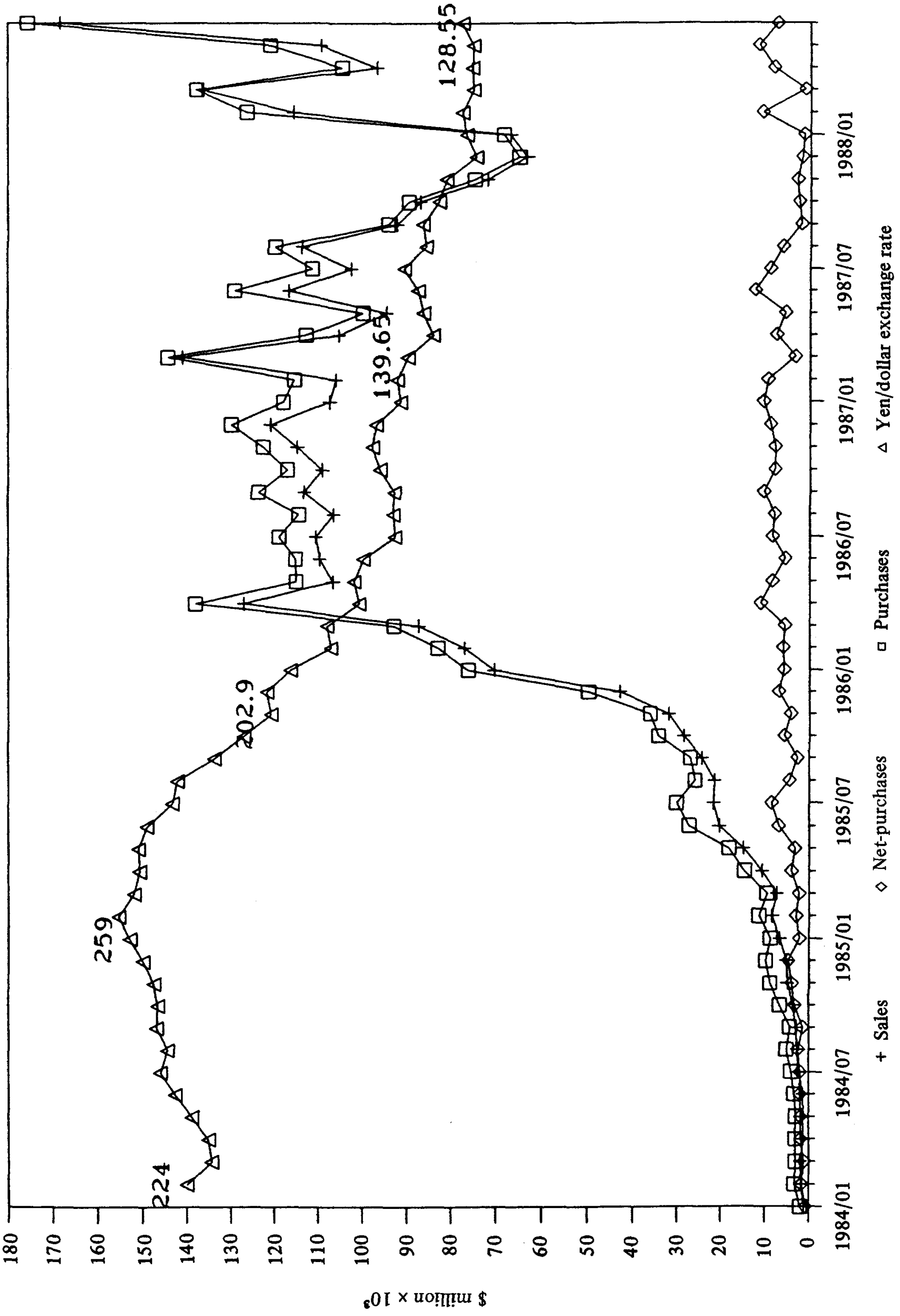


Figure 3. Sales and Purchases of Foreign Bonds by Japanese Investors



phenomenon in the Japanese ten-year bond market, the initiation of trading of futures on bonds in Tokyo (October 1985), and the elimination of the withholding tax on foreign investors imposed by the U.S. Treasury. They are suggestive of substantial increases in the degree of international integration of bond markets and an increased importance of capital flows in the determination of bond yields in Japan. The trading patterns of Japanese investors further suggest that U.S. interest rate movements, in particular, may have had significant effects on the volatility of bond yields in Japan. It is for this reason that shifts in the U.S. term structure of government bonds enter prominently in the empirical analysis of Section IV.

Though the presumption of this analysis is that the levels and volatilities in yields on short- and long-term bonds are inter-related, a specific model of the term structure linking the temporal behavior of short- and long-terms bonds is not imposed on the data. Instead, the econometric relations set forth in Section III are chosen in order to provide a flexible descriptive representation of the distribution of bond yields. Perhaps the primary reason for not conducting the empirical analysis in the context of a formal model of the term structure is that the market structures that underlie many expectations based theories of the term structure were not in place in Japan during the sample period. Specifically, short sales of bonds were tightly restricted in the spot market and there was not an active organized market for borrowing securities.[6] Furthermore, the futures markets, which have existed only since October 1985, were available only for ten and twenty year bonds. Thirdly, as noted above, there was relatively little liquidity in the intermediate-term bond markets. Therefore, the arbitrage trading that links the long and short ends of the yield curve in expectations theories was relatively costly in Japan.

III. Representations of Volatility in Holding Period Yields

Numerous approaches to modeling the conditional variances of asset returns have recently been proposed in the literature. Perhaps the most widely studied specification is the ARCH model originally proposed by Engle (1982), and subsequently applied by Engle, Lilian and Robbins (1987) to the U.S. market and Bomhoff and Schotman (1987) to the U.S. and Japanese bond markets. In these applications, the expected excess holding period returns on long-term bonds are interpreted as risk premia and are assumed to be proportional to the conditional covariance matrix of the vector of returns studied (the ARCH-M model).

Though convenient to implement, the ARCH model of conditional volatility has several limitations as a model of the conditional variances of bond yields. First, the assumed proportionality of risk premia to the conditional covariance matrix of returns is

[6] In June 1989 borrowing-lending markets, similar to the U.S. repurchase markets, were authorized by the Ministry of Finance. As of August 1980, these markets were being shunned by dealers and volume was low. Fees were also sufficiently high to make borrowing not economical.

an implication only of very special models of asset price determination (Glosten, Jagannathan and Runkle 1989). Most dynamic models do not imply a simple linear relation among these moments. Second, the conditional distributions of disturbances in ARCH specifications are typically assumed to be normal. However, there is growing evidence that many asset returns are not well described by a conditional normal distribution. For instance, Gallant, Hseih and Tauchen (1987) show in the context of a semi-nonparametric specification of the distribution of daily exchange rate changes that exchange rates are not adequately described by an ARCH model with conditional normal or conditional t distributions. Furthermore, the findings in Pagan and Hong (1988) suggest that the conditional variance of excess U.S. Treasury bill yields is not a symmetric function of past yield differentials or forecast errors as is assumed in most ARCH specifications.[7] They reject the ARCH model for U.S. Treasury bill returns.

In the light of these findings, a nonparametric specification of the conditional means and variances of the holding period returns on JGBs is adopted. Specifically, I follow Pagan and Ullah (1988) and Pagan and Hong (1988) and adopt Fourier series approximations. The construction of the estimates for the conditional variances proceeds as follows. Let H_t^i denote the i-week holding period return from date t$-$i to date t on a bond with a maturity of n weeks, $i \leq n$. Let x_t denote a vector of variables that is in investors' information set at date t and which generates the conditioning information set for expectations. Though this information set is dated at t, the vector x_t will in general include lagged values of various financial variables. The expected holding period return from t to t+i conditioned on x_t is approximated by fitting a truncated Fourier series representation as in Gallant (1982):

$$E[H_{t+i}^i \mid x_t] = \beta_o + \beta_1' x_t + \beta_2' [x_t x_t'] + \sum_{j=1}^{m} [\sigma_j' \cos(jx_t) + \phi_j' \sin(jx_t)] \tag{1}$$

where $[x_t x_t']$ denotes the vector of distinct elements of the cross-product matrix $x_t x_t'$, $\cos(jx_t)$ denotes the vector of cosine transformations of the elements of jx_t, etc. This approximation extends the commonly used autoregressive representation of expectations by including quadratic and trigonometric terms.

Fourier expansions approximate arbitrarily closely continuous functions on the interval $[0,2\pi]$ as m becomes large.[8] In applying the Fourier methods, the data is scaled so that the sample lies within the interval $(0,2\pi)$. Among the potential problems with using such an approximation—and indeed most nonparametric methods—is that the conditioning information and the variable being forecasted may exhibit growth over time. And growing series will not satisfy the assumption that the transformed series, constructed using a

[7] Nelson (1987) has proposed one possible modification to the standard ARCH model that allows for asymmetric responses of conditional variances to changes in the conditioning variables.

[8] General Fourier series approximations involve sine and cosine functions of linear combinations of the elements of x. In order to keep the size of the parameter space relative to the sample size from becoming large, I consider only sine and cosine functions of the individual elements of x.

fixed transformation function, stay within the approximation interval (here $[0,2\pi]$). The potential importance of this observation for this study is discussed after describing the data used in the analysis.

Having approximated the conditional mean, the forecast errors $\varepsilon_{t+i}=H^i_{t+i}-E[H^i_{t+i}|x_t]$ are formed and then the conditional variance is modeled as

$$\sigma^2_{Ht}(i) = E[\varepsilon^2_{t+i} \mid x_t] = \alpha_0+\alpha_1'x_t+\alpha_2'[x_tx_t'] + \sum_{j=1}^{m} [\Psi_j' \cos(jx_t) + \theta_j' \sin(jx_t)]. \qquad (2)$$

Unlike ARCH representation of conditional variances, lagged values of ε^2_{t+i} do not enter explicitly in x_t in this formulation. However, the ARCH model with a linear conditional mean implies a representation of $\sigma^2_{Ht}(i)$ which can be approximated very closely by (2) with a sufficiently long history of the levels and squared values of H^i_t in x_t. At the same time, (2) is more flexible than ARCH models because of the inclusion of trigonometric terms.

If the slope coefficients in (2) are not all zero, then the disturbance term in (1) is conditionally heteroskedastic. The forecast error $(\varepsilon^2_{t+i}-\sigma^2_{Ht}(i))$ may also be heteroskedastic, so heteroskedastic-consistent standard errors (Hansen 1982) are presented for the parameter estimates. The estimate of the parameter vector $\alpha'=[\alpha_0, \alpha_1, \alpha_2, \Psi'_1, \theta'_1, \ldots, \Psi'_m, \theta'_m]$ for (2) is calculated using a two-step estimation procedure: first the parameter vector $\beta'=[\beta_0, \beta_1, \beta_2, \sigma'_1, \phi'_1, \ldots, \sigma'_m, \phi'_m]$ for (1) is estimated and then the squared values of the estimated residuals are taken to be the dependent variable when fitting the variance equation (2). Notice, however, that the standard errors from the second stage do not need to be adjusted for the use of the first stage estimator of β. This follows from the observation that, under the null hypothesis $E[\varepsilon_{t+i}\partial\varepsilon_{t+i}/\partial\beta]\equiv E[\varepsilon_{t+i}h(x_t)]=0$; see Newey (1984).

Two values of i are considered:1 and 13. The first value corresponds to the case of weekly holding period returns on ten-year bonds, while the second case corresponds to 13 week (quarterly) holding period returns. When $i=1$, the disturbance ε_{t+1} is serially uncorrelated by construction as long as x_t includes a sufficiently long history of H^1_t.[9] In the case of $i=13$, the forecast error ε_{t+13} follows an MA(12) process (again assuming that the history of holding period returns is included in x_t). Therefore, the standard errors of the estimated parameters in both (1) and (2) are also adjusted for the serial correlation of the disturbances.[10]

[9] The projection error $(\varepsilon^2_{t+1}-\sigma^2_{Ht}(1))$ may be serially correlated since the past history of the squared forecast errors is not included in x_t. Standard errors and test statistics were calculated both under the assumption that $(\varepsilon^2_{t+1}-\sigma^2_{Ht}(1))$ is serially uncorrelated and under the assumption this error follows an MA(6) process. The results were qualitatively similar with the tests of exclusion restrictions providing more evidence against the null hypotheses under the MA(6) assumption. The results assuming no serial correlation are presented for the case $i=1$.

[10] The MA(12) error in (1) with $i=13$ can in principle be accommodated using the estimator of the standard error discussed by Hansen (1982) and Hansen and Singleton (1982). In practice, however, this approach lead to a non-positive definite parameter covariance matrix and, therefore, the Newey-West procedure with twelve lags was also applied for this equation.

Turning next to the selection of the data for the empirical analysis, there are several choices for the returns on long-term JGBs. Typically, in studying the term structure, holding period returns on a sequence of bonds with a common fixed maturity are examined. A potential drawback of examining such a series for Japan is that many of the bonds considered will not be benchmark bonds at the time the holding periods are calculated. And, as noted in Section II, there will be relatively small trading volumes in non-benchmark bonds. A series of holding period returns on the actively traded benchmark securities can be constructed. Since a given bond issue remains the benchmark bond for several months, the constant maturity assumption underlying term structure analyses will be violated more severely for this series. Nevertheless, since the benchmark is the most actively traded security, it may more rapidly and accurately reflect domestic and international developments which affect economic activity in Japan. Results for holding period returns on both the benchmark and constant maturity bond series will be examined.

Holding period returns were calculated for the benchmark series from a series of daily observations on closing prices on the Tokyo Stock Exchange (TSE). A comparable series was not available for the constant maturity series. Instead, I constructed a weekly over-the-counter (OTC) series of yields-to-maturity on the bonds with maturities closest to ten years. Then approximate weekly holding period returns on this constant maturity series of bonds were calculated as follows. As in Shiller (1979), Singleton (1980), and Shiller, Campbell and Schoenholtz (1983), the i-week holding period return, H^i_t, on a bond with maturity "in" weeks was approximated by

$$H^i_{t+i} = [R^{in}_t - \bar{\gamma} R^{i(n-1)}_{t+i}]/(1-\bar{\gamma}), \tag{3}$$

where $\bar{\gamma}=(\gamma-\gamma^n)/(1-\gamma^n)$ and γ was set equal to $1/(1+R^*)$ with R^* being the average yield-to-maturity of the bond over the sample period.

In order to assess the potential magnitudes of the approximation errors from using (3) instead of the exact holding period returns, the weekly holding period returns ($i=1$) were calculated for the benchmark series using the price data (exact returns) and the associated yields-to-maturity (equation (3)). The results are displayed in Figure 4. The two series track each other quite closely, suggesting that the approximation errors are in practice small. The approximation errors are likely to be smaller for the constant maturity series, since fixed values of γ and n were used in (3) even though a benchmark bond may have remained the benchmark for nearly a year.

The first part of Table 1 displays the issue numbers and the periods during which each issue served as the benchmark for the sample period March 14, 1986 through May 20, 1988. The second part of this table displays the corresponding information for the bond with maturity closest to ten years (the most recently issued ten-year JGB). The constant maturity yields are end of week values. Since the bond markets in Japan were open on some Saturdays, the end of the week could be at the close of the market on

Figure 4. Seven Day Return – Returns from Yields and Prices

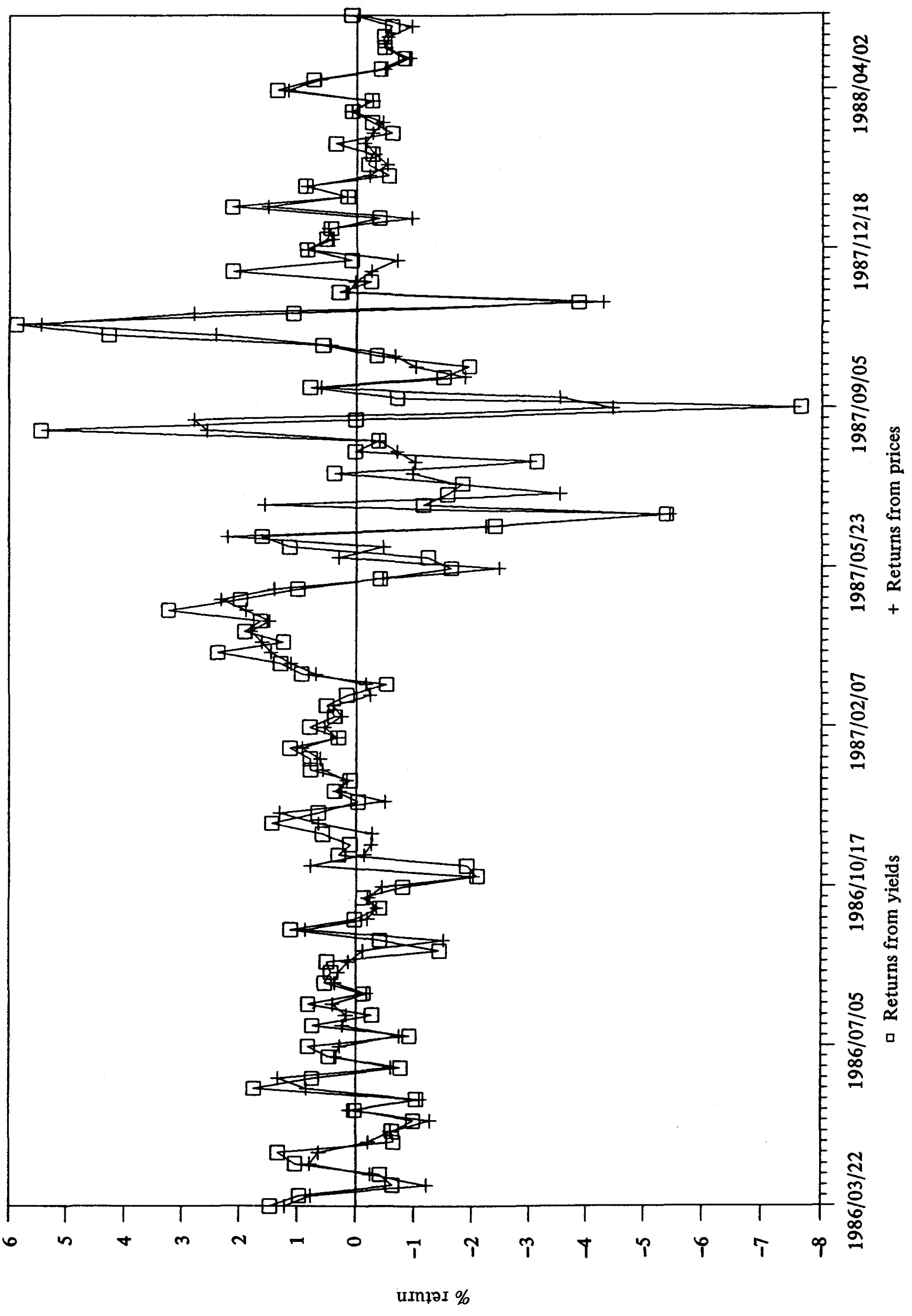


Table 1. Composition of JGB Series

Benchmark Series

March 14, 1986 – October 25, 1986	Issue 78
November 1, 1986 – November 28, 1987	Issue 89
December 5, 1987 – May 20, 1988	Issue 105

Fixed Maturity Series

March 14, 1986 – May 9, 1986	Issue 80
May 17, 1986 – August 15, 1986	Issue 89
August 23, 1986 – November 29, 1986	Issue 90
December 6, 1986 – December 27, 1986	Issue 92
January 9, 1986 – March 7, 1987	Issue 93
March 13, 1987 – May 2, 1987	Issue 97
May 8, 1987 – May 30, 1987	Issue 99
June 6, 1987 – August 7, 1987	Issue 100
August 14, 1987 – September 5, 1987	Issue 102
September 11, 1987 – October 9, 1987	Issue 103
October 16, 1987 – November 28, 1987	Issue 104
December 5, 1987 – January 30, 1988	Issue 105
February 6, 1988 – February 27, 1988	Issue 106
March 5, 1988 – April 8, 1988	Issue 108
April 15, 1988 – May 20, 1988	Issue 109

Table 2. Descriptive Statistics for Weekly Yields
March 14, 1986 – May 20, 1988

Series	Mean	S.D.	Autocorrelations						γ
			1	2	3	4	5	6	
OTC	.00094	.00012	.938	.844	.765	.693	.610	.515	
OTF	.00100	.00011	.925	.845	.766	.684	.584	.479	
HOTC	.00119	.01658	.287	–.117	–.065	.089	.105	.061	.997
HOTF	.00107	.01652	.051	–.003	.020	.120	.043	.080	.997
HTSE	.00109	.01460	.201	–.059	.066	.020	.019	–.068	
HTSEW	.00022	.01463	.205	–.055	.061	.019	.023	–.065	

Friday or Saturday. Comparisons of weekly holding period returns on the benchmark bonds calculated using TSE data and either end of week or Friday closing prices suggest that the time series properties of these two return series were very similar. That is, the small amount of additional variation in maturity induced by occasional Saturday trading had small effects on the yields constructed. Subsequent results for the benchmark series are based on closing prices for Friday, while the constant maturity results are based on the end of week yields.[11]

Descriptive statistics for the various series are presented in Table 2. The weekly OTC yields-to-maturity on the benchmark series are denoted OTC, while the yields-to-maturity on the constant maturity OTC series is denoted OTF. Comparing the statistics for OTC with those for OTF, one sees that the average return on OTC was smaller than the average return on OTF. This difference is the manifestation of the liquidity premium associated with the relatively low volume of trade in OTF. The autocorrelations of the series are comparable.

The remaining rows in Table 2 display descriptive statistics for the holding period returns. When $i=1$, the implied values of $\bar{\gamma}$ for the OTC and OTF series is .997. Therefore, construction of the approximate holding period returns using (3) amounts to taking a quasi-difference of the yields with a filter that has a root very close to the unit circle. As a consequence, the approximate holding period returns, HOTC and HOTF, exhibit substantially less autocorrelation than the underlying yields-to-maturity. Indeed, in the case of HOTF, the autocorrelations are all near zero which suggests that the conditional distributions of HOTC and HOTF may be very different. The last two rows display statistics for the weekly holding periods constructed from daily TSE benchmark prices with (HTSE) and without (HTSEW) including the coupon payments. The standard deviations and autocorrelations of these series are comparable to those of HOTC as might be expected from the patterns in Figure 4.

A notable feature of Figure 4 is the substantial increase in the volatilities of the benchmark holding period returns during 1987. The increased volatility between January and April of 1987 may have been attributable to the substantial increase in the trading volume in the benchmark JGB (Figure 1). Concurrent with this high trading volume, there was a substantial decline in the yield on the benchmark series to the point where the benchmark yield was approximately equal to the Bank of Japan's discount rate.

The largest swings in returns were after this period of high turnover in JGB markets, however. There was a substantial appreciation of the yen against the dollar between August and December 1987. The yen/dollar exchange rate reached 152 in late August and then fell to 122 by late December. During most of this period Japanese investors were net sellers of foreign (and in particular U.S.) bonds, and the turnover of foreign

[11] Of course, for some weeks there were holidays on Friday or Saturday in which case the yield was calculated at the close of the day closest to the end of the week.

bonds fell substantially (Figure 3). Moreover in September the long-term JGB yields rose above 6% (they were 4.11% in June 1987), while long- and short-term U.S. Treasury bond yields rose sharply through the third quarter of 1987 and then fell equally sharply at the end of 1987. Thus, the relatively large fluctuations in holding period yields during the second half of 1987 coincided with changes in exchange rates and yield differentials between Japan and the U.S., and associated international capital flows.[12]

Therefore, the following variables are included in the conditioning information set used to estimate the conditional variances. First, two yields from the U.S. Treasury bond market were included: the yield on three-month Treasury bills and the constant maturity yield on ten-year Treasury bonds. Both series were constructed from the daily series compiled by the Federal Reserve. Also, the yen/dollar exchange rate as recorded in the Tokyo foreign exchange market was included. All of these series were recorded at the close of trading on the day before the yields in the Japanese bond market were recorded to assure that they were known to market participants at the time OTC and TSE prices were sampled.

Finally, the yield on a yen-denominated short-term bond is included. In selecting this yield, consideration was given to the fact that the short-term government bond markets were not as well developed in Japan as in some other countries. The treasury bill market is in its infancy. Issues of short-term government debt have been infrequent and the stock of outstanding bills at any one time is small compared to, say, the U.S. Treasury bill market. In addition, income from government bills in Japan is subject to a withholding tax for residents and nonresidents.[13] In contrast, there are active secondary markets for large denomination Euroyen deposits. The yields on these deposits are studied in Section IV.

IV. Empirical Results

Before estimating the mean and variance relations (1) and (2), the data were transformed to induce stationarity. The autocorrelations of the holding period returns displayed in Table 2 suggest that these returns are stationary. The U.S. Treasury bill data was transformed by time differencing the three-month Treasury bill yield (DTB3) and taking the contemporaneous difference between the yield on the ten-year bond and the three-month yield (SPD). The latter spread variable is a measure of the shape of the term structure of Treasury bonds. Finally, the yen/dollar exchange rate (DY$) and the yield on one-week Euroyen deposits were first-differenced (DEURO). These transformed series

[12] The financial market uncertainty during this period was compounded by the stock market crash in October 1987.

[13] Because of the lack of depth in the Japanese government bill market, open market operations in the conventional sense of this term is not a primary tool of monetary policy in Japan. The Bank of Japan trades in security markets, but primarily in the commercial bill market (Suzuki 1987).

were assumed to be free of significant trends over the sample period.

The sample period examined was March 14, 1986 through May 20, 1988. The choice of the beginning date reflects the large increase in trading volume in foreign markets by Japanese investors up through the early part of 1986 (see Figure 3). In addition, the beginning date is after the initiation of trading in long-term bond futures in October 1985 and the notable intervention by the Bank of Japan in domestic financial markets in late 1985.

Initially, only the linear and quadratic terms were included in (1) and (2); the sine and cosine terms were omitted. Since the transformation of the data to a range of $[0,2\pi]$ amounts to a simple rescaling of the regressors, the test results for this quadratic approximation are identical to those that would be obtained from estimating a quadratic representation of the untransformed data. Thus, the results from these regressions permit a preliminary assessment of the gains in fit from including quadratic terms in the commonly studied linear forecasting models for holding period yields.

The results for weekly holding period returns on the benchmark (HTSE) series are displayed in Table 3. Column one displays the results for the mean equation (1) and column two displays the estimates for the conditional variance (2). The variables H2, EURO2, DY$2, DTB2, and SPD2 represent the squared values of HTSE, DEURO, DY$, DTB, and SPD, respectively. The subscripts indicate the number of time periods the variable is lagged in the regression (e.g., $HTSE_1$ is HTSE lagged one week). The numbers in parentheses are the estimated values of the asymptotic standard errors for the parameters and marginal significance levels for the chi-square tests of exclusion restrictions. A "*" (**) adjacent to a standard error of an estimate indicates the null hypothesis that the coefficient is zero is rejected at the 5% (1%) significance level.

There was clearly significant variation in the conditional mean of HTSE during this sample period. Furthermore, the results suggest that the inclusion of quadratic terms increases significantly the explanatory power of the regression: the estimated coefficients on $H2_2$ and $DTB2_2$ are significant at the 1% level based on the asymptotic standard errors. Chi-square statistics for the joint hypothesis that the linear and quadratic terms for each of the variables all have zero coefficients are presented at the bottom of Table 3. The subscripts E, Y, T, and S indicate DEURO, DY$, DTB, and SPD, respectively. The joint hypothesis for DTB is rejected at conventional significance levels.

There is also evidence of substantial time variation in the conditional variance of the holding period return on the benchmark series. The estimates displayed in the second column of Table 3 suggest that changes in the U.S. Treasury bill rate, the Euroyen deposit rate, and long-short U.S. Treasury yield spread predicted changes in $\sigma^2_{Ht}(1)$. Thus variables which, according to conventional significance levels, do not predict changes in the conditional mean may impact significantly on the conditional variance. Compare, for example, the results in the first and second columns of Table 3 for SPD and DEURO. In this sense, the sources of risk underlying trading in long-term bonds may not be reliably

Table 3. Estimates of Quadratic Regressions for Benchmark Weekly Holding Period Returns
March 14, 1986 – May 20, 1988

	HTSE	ϵ_t^2
$HTSE_1$	−.5243 (.398)	−.6337 (.624)
$HTSE_2$	1.651 (.288)**	.1786 (.407)
$HTSE_3$	.2014 (.082)	−.0034 (.142)
$HTSE_4$	.1761 (.075)*	−.0403 (.092)
$H2_1$	.1444 (.088)	.1281 (.136)
$H2_2$	−.3204 (.057)**	.0295 (.095)
$DEURO_1$	−.2998 (.366)	−.0322 (.493)
$DEURO_2$	.1066 (.273)	.8829 (.431)
$DEURO_3$	.0326 (.094)	−.1397 (.139)
$EURO2_1$	.0543 (.049)	−.0354 (.066)
$EURO2_2$	−.0137 (.058)	−.1804 (.082)
$DY\$_1$	−.1384 (.332)	−.4551 (.376)
$DY\$_2$	.0498 (.347)	−.6429 (.479)
$DY\$_3$	.0871 (.075)	.0513 (.091)
$DY\$2_1$	.0090 (.048)	.0705 (.050)
$DY\$2_2$	.0086 (.050)	.1033 (.069)
DTB_1	−1.534 (.314)**	.2166 (.948)
DTB_2	−.4085 (.539)	1.537 (.776)*
DTB_3	.0312 (.146)	.6475 (.217)*
$DTB2_1$	.1888 (.044)**	−.0210 (.060)
$DTB2_2$	.0414 (.057)	−.1789 (.083)
SPD_1	.3830 (.418)	.5793 (.483)
SPD_2	−.4519 (.485)	.1050 (.638)
SPD_3	−.0075 (.192)	.0393 (.296)
$SPD2_1$	−.0396 (.061)	−.1198 (.075)
$SPD2_2$	.0458 (.070)	.0202 (.089)
R^2	.482	.322
χ_E^2 (5)	6.888 (.229)	15.65 (.008)
χ_Y^2 (5)	2.644 (.755)	5.606 (.346)
χ_T^2 (5)	64.14 (.000)	15.58 (.008)
χ_S^2 (5)	1.820 (.873)	11.90 (.036)

Note: * (**) indicates the null hypothesis that the coefficient is zero is rejected at 5% (1%) significance level.

identified from analyses of correlations between the level of HTSE and other macroeconomic variables.

A potentially important caveat regarding interpretations of estimates of conditional volatility based on the projection equation (2) is that they will in general be sensitive to the specification of the conditional mean (1), since the *a priori* specification of the conditional mean identifies the conditional variance. If, for instance, the conditional mean is a nonlinear, nonquadratic function of the conditioning information, then fitting a quadratic function for the mean may lead to an overestimate of the degree of variation in the conditional variance. On the other hand, even if (1) is correctly specified, quadratic versions of (2) may underestimate the degree of conditional variation in returns if the conditional variance is a nonquadratic function of the conditioning information.

To explore the robustness of the findings in Table 3 to alternative functional forms for the conditional moments, a Fourier series approximation including sine and cosine terms was estimated. Specifically, m was set to 3 in (1) and (2) for $HTSE_1$ and $DY\$_1$, while m was set to 2 for DTB_1; only trigonometric functions of the first lag of these variables were added.[14] Chi-square statistics for various exclusion restrictions on these regressors are presented in Table 4. Each column represents a different specification of the conditioning information. Runs with trigonometric terms for all of the variables were not attempted, because of the large number of regressors relative to the sample size for such models. Additionally, DEURO and SPD, as well as their squared values, were omitted from all of the regressions. A larger percentage of the variation in HTSE is explained by all of the approximations in Table 4 compared to the quadratic approximation underlying Table 3. And the coefficients of determination for the ε_t^2 equations are correspondingly smaller in Table 4 relative to Table 3. Fourier terms enter significantly in all of the projection equations.

A more direct comparison of the conditional variances implied by the quadratic and Fourier approximations is provided by Figures 5 and 6 which display the actual squared residuals and their respective fitted values from the quadratic model (Model 3) and the Fourier model in the first column of Table 4 (Model 4.1). Overall, the patterns of squared residuals, $\hat{\varepsilon}_t^2$, are similar across the two approximations. The relatively smaller values of $\hat{\varepsilon}_t^2$ during 1987 and 1988 in Figure 6 compared to Figure 5 are a consequence of the additional predictive power of the Fourier terms for the conditional mean of HTSE. Similarly, the estimated conditional variances displayed in Figures 5 and 6 follow qualitatively similar time paths.[15] Together, Figures 5 and 6 suggest that the inclusion of Fourier terms in the approximation does not alter the qualitative features of the estimated conditional variances.

Interestingly, the largest conditional variances in Figure 6 occur around April 1987

[14] Attempts to estimate the models with m=3 for DTB often failed because of multicollinearity in the series.

[15] The conditional variance equation (2) does not constrain the estimated conditional variances to be positive and a few of the point estimates did take on small negative values in the sample.

Table 4. Chi-square Statistics for Fourier Series Approximation Benchmark Series
March 14, 1986 – May 20, 1988

		HTSE	
$\chi^2_{HF}(6)$	33.96 (.000)	.....	11.60 (.071)
$\chi^2_{Y}(5)$	21.29 (.000)	30.59 (.000)	6.842 (.232)
$\chi^2_{YF}(6)$	56.78 (.000)	33.22 (.000)	.....
$\chi^2_{T}(5)$	59.98 (.000)	54.68 (.000)	17.02 (.004)
$\chi^2_{TF}(4)$	.....	6.531 (.163)	1.973 (.741)
R^2	.619	.524	.535
		Squared residual	
$\chi^2_{HF}(6)$	17.07 (.009)	.....	11.06 (.087)
$\chi^2_{Y}(4)$	18.17 (.003)	5.457 (.363)	7.357 (.195)
$\chi^2_{YF}(6)$	16.29 (.012)	9.154 (.165)	.....
$\chi^2_{T}(4)$	14.59 (.012)	11.80 (.038)	7.295 (.200)
$\chi^2_{TF}(4)$	.....	7.090 (.131)	9.444 (.051)
R^2	.314	.295	.285

HF = Fourier terms for HTSE
Y = Linear and quadratic terms for DY$
YF = Fourier terms for DY$
T = Linear and quadratic terms for DTB
TF = Fourier terms for DTB
Marginal significance levels of the test statistics are given in parentheses.

when there was a large increase in the trading volume in government bonds. More precisely, volatility rose along with volume, fell near the peak of the high turnover ratios at the end of April 1987, and then conditional volatility increased as volume fell to more normal levels. Moreover, this period is also characterized by the largest deviations between the actual and fitted values of $\hat{\varepsilon}_t^2$. That is, the conditional variances, as measured by (2), understated by substantial margins the actual squared forecast errors for returns.

Other periods of relatively large conditional variances for the holding period returns were July-August 1987 and December 1987 (especially for the quadratic approximation). As noted previously, the yen/dollar exchange rate peaked in late August 1987 and fell substantially until December 1987. Furthermore, July 1987 was the beginning of a significant decline in the trading of foreign bonds by Japanese investors. This decline continued until November 1987 and then the trading of foreign bonds by Japanese increased substantially (Figure 3). These observations are consistent with the view that

Figure 5. Squared Deviation and their Fitted Values: Model 3

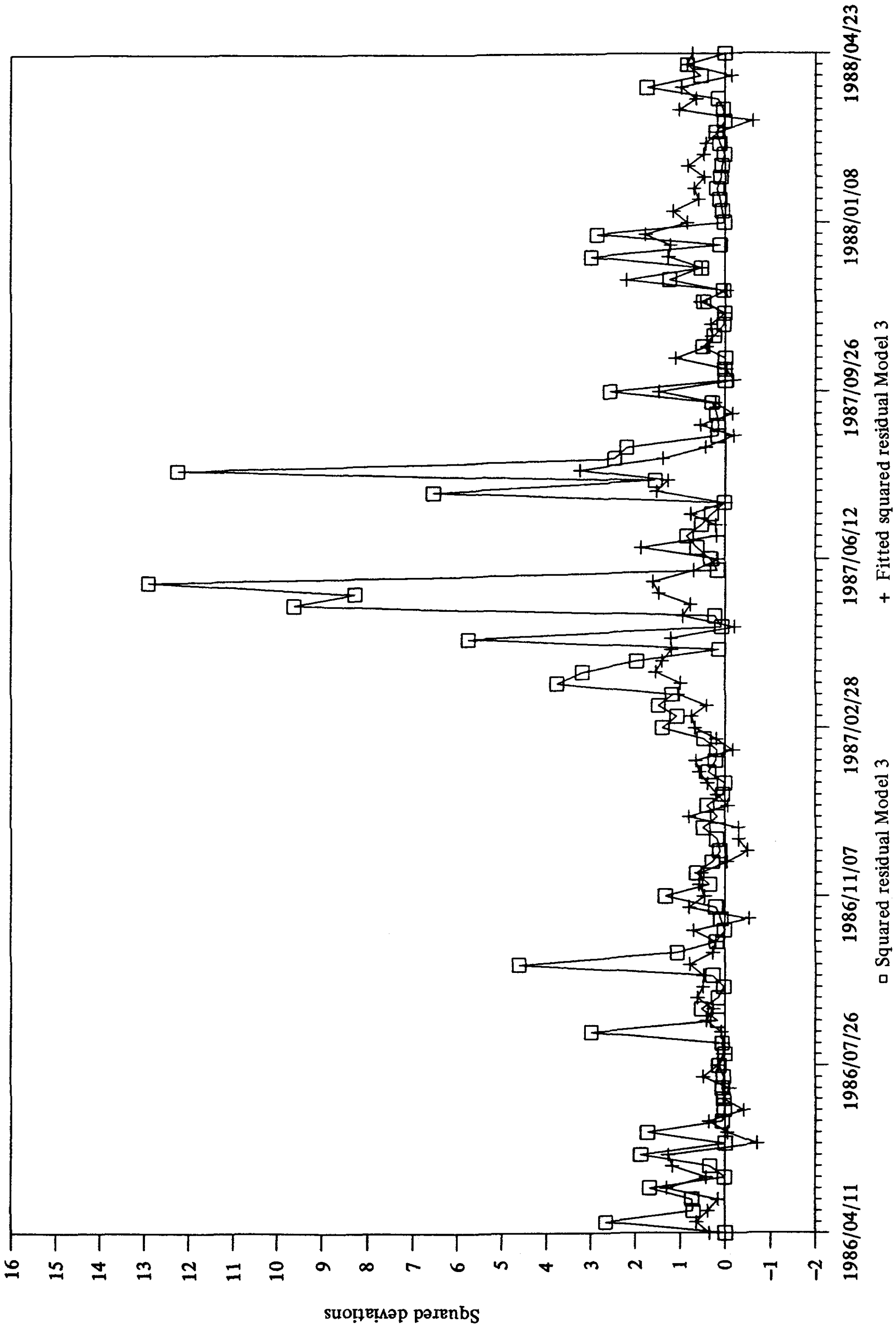
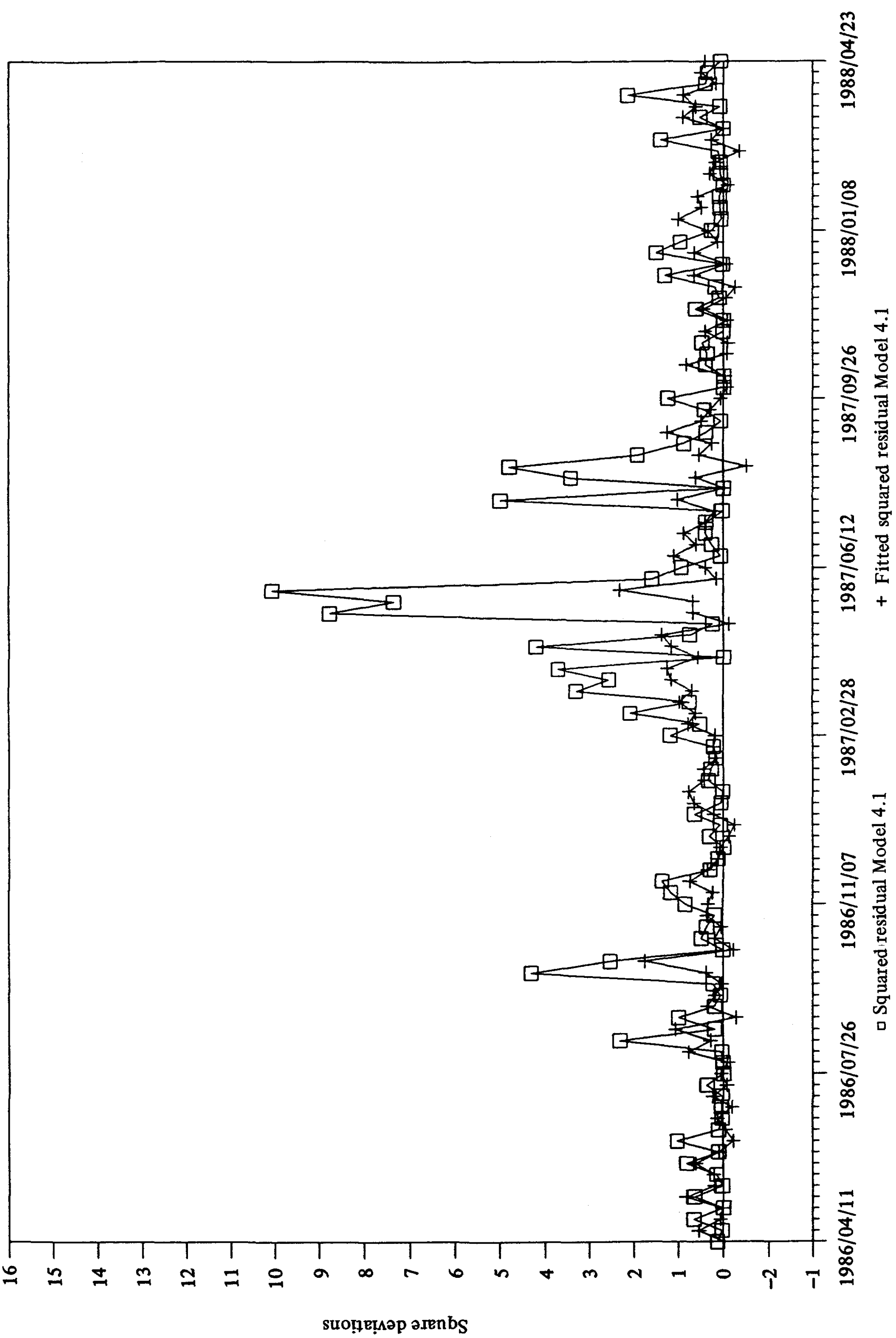


Figure 6. Squared Deviation and their Fitted Values: Model 4.1



exchange rate changes and capital flows to and from Japan were associated with increased *ex ante* volatility in the Japanese bond markets. The regression results in Tables 3 and 4 suggest in addition that changes in short-term interest rates in the U.S. were highly correlated with these capital flows and volatility changes.

Estimates of the quadratic approximations to the conditional moments for quarterly (13 week) returns are displayed in Table 5. The explanatory variables were lagged at least 13 periods to account for the 13 week forecast horizon associated with the expectations in (1) and (2). Also, the weekly Euroyen deposit rate was replaced with the 13 week deposit rate in order to match the length of the holding period for the benchmark JGB. For this longer forecast horizon, all of the conditioning variables enter the mean approximation significantly at conventional significance levels except for the Euroyen deposit rate. These findings are consistent with previous monthly studies in which exchange rates and U.S. interest rates were statistically significant explanators of yields on benchmark JGBs.

The results for the estimated conditional variances (last column of Table 5) are similar to those obtained for weekly holding period returns. The Euroyen deposit rate and the U.S. Treasury yield spread are again significant explanators of the squared forecast errors, while the coefficients on changes in the three-month U.S. Treasury bill yield are somewhat less significant as a group in the 13 week forecast equation. The diminished role for DTB relative to SPD in the conditional variance equations for 13 week holding periods may reflect different investment considerations over 13 versus 1 week holding periods.

For comparison, the estimates of the quadratic approximations to the moments of the constant maturity, holding period returns are displayed in Table 6. The holding period return (HOTF) was calculated using the OTC data and the approximation (3), and squared values of HOTF are denoted by HF2. As with HTSE, there is significant variation in the conditional mean of HOTF, and this variation is associated primarily with past changes in HOTF and the U.S. Treasury bill yield.

While the test results for the estimated conditional means of HTSE and HOTF are similar, the results for the (quadratic approximate) conditional variances are notably different.[16] The relatively small marginal significance level for the block of coefficients on SPD suggest that volatility of the constant maturity, holding period yields does change sympathetically with the shape of the term structure of interest rates in the U.S. But none of the other financial variables seem to have significant explanatory power for the conditional variance of HOTF. This result is consistent with the observations in Section II about the institutional buy-and-hold strategies for off-the-run bonds and the absence of

[16] This finding appears to be inconsistent with the ARCH in mean specification of expected returns adopted by Bomhoff and Schotman (1987) for the Japanese bond market. It is interesting to note that, using very different data and a different sample period, they found that the ARCH effect on the conditional mean was insignificant. For an ARCH in mean model this implies that the risk premium on the bond is constant, but the evidence for reported here for constant maturity bond series does not support this conclusion.

Table 5. Estimates of Quadratic Regressions for Benchmark Quarterly Holding Period Returns
March 14, 1986 – May 20, 1988

	HTSE	ϵ_t^2
$HTSE_{13}$	.6799 (.343)*	–.6254 (.426)
$HTSE_{14}$	.3326 (.349)	.1599 (.480)
$HTSE_{15}$	.1267 (.204)	.0729 (.304)
$HTSE_{16}$	–.7161 (.219)**	–.1822 (.345)
$H2_{13}$	–.0380 (.051)	.1240 (.082)
$H2_{14}$	–.1015 (.034)**	–.0225 (.086)
$DEURO_{13}$	–.2731 (.284)	.7423 (.312)*
$DEURO_{14}$	–.0178 (.289)	.0479 (.346)
$DEURO_{15}$	–.1238 (.113)	.2197 (.141)
$EURO2_{13}$	.0278 (.042)	–.1133 (.060)
$EURO2_{14}$	–.5045 (.041)	–.0317 (.053)
$DY\$_{13}$	.0750 (.357)	.1157 (.317)
$DY\$_{14}$	.7761 (.423)	.0462 (.543)
$DY\$_{15}$	.1259 (.149)	.0459 (.226)
$DY\$2_{13}$	.0075 (.045)	–.1019 (.035)
$DY\$2_{14}$	–.0834 (.054)	–.0345 (.068)
DTB_{13}	.2399 (.293)	.5017 (.296)
DTB_{14}	.9531 (.538)	.4744 (.425)
DTB_{15}	.5354 (.140)**	–.2949 (.190)
$DTB2_{13}$	.0419 (.066)	–.0994 (.048)*
$DTB2_{14}$	–.0414 (.088)	–.0969 (.067)
SPD_{13}	–.9756 (.515)	–2.021 (.701)**
SPD_{14}	–.4714 (.508)	.4785 (.712)
SPD_{15}	–.0708 (.195)	.5664 (.230)*
$SPD2_{13}$	.1648 (.072)*	.2647 (.093)**
$SPD2_{14}$	.0827 (.086)	–.1447 (.104)
R^2	.623	.364
$\chi_E^2(5)$	2.332 (.801)	11.75 (.038)
$\chi_Y^2(5)$	15.17 (.009)	5.846 (.321)
$\chi_T^2(5)$	29.57 (.000)	8.185 (.146)
$\chi_S^2(5)$	32.65 (.000)	22.05 (.000)

Note: * (**) indicates the null hypothesis that the coefficient is zero is rejected at 5% (1%) significance level.

Table 6. Estimates of Quadratic Regressions for Constant Maturity Weekly Holding Period Returns
March 14, 1986 – May 20, 1988

	HOTF	ϵ_t^2	HOTF-EURO
$HOTF_1$	.2980 (.256)	–.0480 (.178)	.2929 (.256)
$HOTF_2$	.8436 (.303)**	.2162 (.234)	.8388 (.303)**
$HOTF_3$	.0599 (.090)	–.0626 (.048)	.0585 (.090)
$HOTF_4$	.2206 (.117)	.0774 (.090)	.2201 (.118)
$HF2_1$	–.0848 (.039)**	–.0020 (.027)	–.0840 (.039)**
$HF2_2$	–.1507 (.050)**	–.0213 (.040)	–.1502 (.050)**
$DEURO_1$	–.7164 (.426)	–.0544 (.194)	–.7178 (.426)
$DEURO_2$	.0704 (.426)	.3288 (.256)	.0686 (.332)
$DEURO_3$	–.1974 (.119)	.0043 (.087)	–.2001 (.119)
$EURO2_1$	.8312 (.062)	.0029 (.030)	.0823 (.062)
$EURO2_2$	–.0519 (.067)	–.0596 (.045)	–.0524 (.067)
$DY\$_1$	–.2163 (.271)	–.0560 (.117)	–.2162 (.272)
$DY\$_2$	.2255 (.300)	–.1060 (.142)	.2242 (.300)
$DY\$_3$	–.0011 (.083)	–.0538 (.057)	–.0008 (.083)
$DY\$2_1$	.0077 (.044)	–.0484 (.022)	.0075 (.044)
$DY\$2_2$	–.0204 (.044)	.0108 (.020)	–.0203 (.044)
DTB_1	–1.390 (.258)**	.0411 (.158)	–1.395 (.258)**
DTB_2	–.6909 (.338)*	.1998 (.234)	–.6943 (.229)**
DTB_3	.1961 (.177)	.0704 (.137)	.1964 (.177)
$DTB2_1$	.1852 (.050)**	–.0438 (.040)	.1859 (.050)**
$DTB2_2$	.0789 (.045)	–.0167 (.027)	.0792 (.045)
SPD_1	–.0641 (.512)	.1716 (.223)	–.0641 (.513)
SPD_2	.0204 (.729)	.8994 (.741)	.0234 (.730)
SPD_3	–.3767 (.320)	–.4159 (.324)	–.3765 (.320)
$SPD2_1$	.0837 (.090)	–.0970 (.041)*	.0838 (.090)
$SPD2_2$	–.0129 (.090)	.0016 (.051)	–.0133 (.090)
R^2	.343	.266	.343
$\chi_E^2(5)$	5.420 (.367)	4.307 (.506)	5.561 (.351)
$\chi_Y^2(5)$	5.587 (.348)	3.347 (.647)	5.585 (.349)
$\chi_T^2(5)$	35.68 (.000)	8.871 (.114)	35.90 (.000)
$\chi_S^2(5)$	7.475 (.188)	10.39 (.065)	7.473 (.188)

Note: * (**) indicates the null hypothesis that the coefficient is zero is rejected at 5% (1%) significance level.

large trading volumes for these securities. Evidently, neither the level or volatility of the constant-maturity yields are as forecastable from near-term changes in exchange rates or short-term bond yields as the benchmark yields.

Though the focus of this analysis is not on the term structure of interest rates in Japan, some insights into the validity of the expectations theories of the term structure can be gained by examining the corresponding results for the excess holding period returns on long-term bonds. The expected difference between the (constant maturity) holding period return H_t^1 and the Euroyen one-week deposit rate $EURO_t$ is often interpreted as the risk premium from holding long-term bonds. Evidence of time variation in this difference is inconsistent with some expectations theories of the term structure of interest rates.

To test whether there was significant time variation in the risk premium, the excess returns (H_t^1–$EURO_t$) for HOTF was regressed on the same information used for the quadratic approximations.[17] The results are reported in the third column of Table 6. The point estimates indicate that there is significant time variation in the *ex ante* risk premium. Equally interesting is the close similarity between the parameter estimates in the first and third columns within Table 6. This similarity suggests that there is relatively little variation in EURO compared to variation in HOTF. Put differently, the comovements between expected holding period returns and the Euroyen rate are much too weak to be consistent with at least simple versions of the expectations theory of the term structure. This finding is perhaps not unexpected for the reasons set forth in Sections II and III.

V. Concluding Remarks

This paper has investigated the time series properties of the conditional variances of holding period returns on Japanese government bonds using weekly data on 1 and 13 week holding period returns for the period March 1986 through May 1988. The findings suggest that the conditional mean of the weekly holding period returns on the benchmark bond series is strongly correlated with changes in the three-month U.S. Treasury bill rate. The conditional variances of the holding period returns were also correlated significantly with changes in Euroyen and U.S. Treasury bill rates and the U.S. Treasury bond long-short yields spread. While the Fourier terms in the approximations to the conditional variances often entered significantly as groups, the qualitative patterns of the implied fitted conditional variances were similar to the fitted variances obtained using quadratic approximations.

The large increase in the turnover of foreign bonds by Japanese investors beginning

[17] By the law of iterated projections, the projection of (H_t^1–$EURO_t$) onto a set of regressors x_{t-1} gives the same results as projecting the *ex ante* risk premium onto x_t.

in early 1986 and the concentration of trading by Japanese investors, outside of Japan, in U.S. bond markets suggests a possible explanation for the significant correlations between yields changes in the U.S. and Japan. Namely, to the extent that the large turnover ratios in their portfolios of U.S. bonds are associated with trading in the U.S. Treasury bill markets, the significant correlations may reflect portfolio reallocations between long-term JGBs and Treasury bills over short investment horizons. Over a 13 week investment horizon, the conditional mean of the holding period returns were also significantly correlated with yen/dollar exchange rate changes and the U.S. Treasury long-short spread. Interestingly, for the sample period examined, the largest values of the estimated conditional variances occurred during 1987 and about the same time as substantial changes in trading volumes in Japanese and U.S. government bond markets by Japanese investors.

Overall, the results for the benchmark series suggest that there is a significant predictable component to the variation in the holding period returns. Since it is the conditional variances and covariances of returns which underlie optimal dynamic portfolio decisions, these findings also suggest the statistical representations of conditional second moments examined here may be useful for formulating optimal hedging strategies for bond portfolios that include JGBs.

More generally, the results from this study, together with previous results by the Bank of Japan (1986, 1988), highlight the increased influence of foreign economic developments on the time series properties of JGB yields that has occurred along with the deregulation of financial markets during the 1980s in Japan. As the process of deregulation continues in Japan and Europe, the influence of external factors on volatility is likely to remain significant and may well increase in importance.

The corresponding results for the holding period returns on the relatively thinly traded constant maturity bond series were very different. Only the U.S. long-short yield spread entered the conditional variance equation significantly at the ten percent significance level. This result may be attributable to the institutional influences which seem to underlie the relatively low trading volume in non-benchmark JGBs. As trading activity increases and the costs of shorting on and off-the-run securities decline, the constant maturity yield series should become increasingly correlated with external and internal economic developments. At that time, the observations regarding hedging the risk of fluctuations in benchmark yields may apply with equal force to the risk of changes in yields on non-benchmark securities.

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