The Chinese Interbank Repo Market: An Analysis of Term Premiums

Longzhen Fan and Chu Zhang*

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*Fudan University and The Hong Kong University of Science and Technology (HKUST).
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Abstract

Due to the lack of short-term government bills, the interbank repo market in China has been providing the best information about market-driven short-term interest rates since its inception. In this paper, we examine the behavior of the repo rates of various terms and their term premiums (i.e., the deviations from the pure expectation hypothesis). The work in this paper supplements the study by Longstaff (2000), which reports support for the pure expectation hypothesis over the short range of the term structure using repo data from the US. While we find that the hypothesis is statistically rejected, the term premiums are economically small.

JEL Classification: G12

Key words: repo market, term premiums, the pure expectation hypothesis

1. Introduction

The last twenty years witnessed the fundamental changes in China as it shifted from a centrally planned economy to a market-oriented one. As numerous state-owned enterprises were privatized and new private firms emerged, productivity in China soared to an unprecedented level and the country's GDP growth rate led the world for about a decade. Imports and exports grew exponentially as the country became the workshop of the world. Along with the development of manufacturing sector, the financial markets in China also boomed. The establishment of two national securities exchanges in 1990 marked an important milestone in the development. The admission of China into the World Trade Organization in early 2000s prompted drastic reforms to its banking sector.

In this paper, we study the interest rate behavior in the Chinese interbank repo market. Although the central government has a tight control of the base interest rate as in many industrialized economies, the interest rates in the Chinese interbank market do fluctuate according to the conditions of demand and supply as billions of Renminbi, the Chinese currency, change hands on the daily basis. We focus on the interbank repo market, which is a market for short-term borrowing and lending using high-quality securities as collateral. Our question is how interest rates of various maturities at a given time reflect the expectations of the market participants about future interest rate changes. Following the literature, we examine this question in the context of testing the so-called pure expectation hypothesis. This hypothesis basically states that the term structure of interest rates at any given time is set such that the expected return on rolling over short-term riskfree investments is the same as the corresponding long-term riskfree rate. Although the hypothesis has been tested repeatedly and rejected frequently in many markets, the interest in re-examining it has never abated. Much has been learned about the determination of interest rates in the process. Term premiums, i.e., longer-term rates in excess of what the pure expectation hypothesis predicts, and their determinants have been the most important concept in decisions related to fixedincome investments. A particularly relevant work by Longstaff (2000) documents that the pure expectation hypothesis is supported in the US repo market, in contrast with the findings based on the US Treasury market.

Our results show that term premiums in the Chinese repo market are positive and increasing with the term. The pure expectation hypothesis is statistically rejected. We relate the term premium to the term repo rates themselves, to term spreads, and to the volatility of the short-term repo rate. All these variables contribute to the determination of term premiums. The magnitude of the term premiums, however, is economically small relative to that found in the US Treasury market and elsewhere. The finding in Longstaff (2000) that the US repo rates conform to the pure expectation hypothesis, though not strictly observed in the Chinese interbank repo market, indeed has some merit for repo markets outside the US.

The paper is organized as follows. Section 2 briefly describe the Chinese interbank repo market. Section 3 discusses the unconditional term premiums and unconditional tests of the pure expectation hypothesis. Section 4 reports results of the conditional test of the pure expectation hypothesis and the determinants of conditional term premiums. Section 5 conducts a simulation study to deal with some econometric problems typically found in the studies of interest rates. The last section concludes the paper.

2. The Chinese Interbank Repo Market

The Chinese Treasury (hereafter Treasury) secondary market officially began in 1990 when the securities exchanges were established. The trading activities, however, have been limited. Unlike the case of the US, there are no regular cycles in Chinese Treasury issues. While the total number of Treasury bonds is now becoming nontrivial, most bonds have long terms. There have been very few near-maturity bonds in the last five years. As a result, there are no market-driven short-term rates available from the

Treasury market on a continual basis.

To study the short-term interest rate behavior of the Chinese market, we use the interbank repo market. A repo transaction is a pair of spot and forward security transactions in form, but borrowing and lending with security as collateral in substance. In China, repo transactions are a recent phenomenon. There are two major repo markets. One is the market for exchange-traded repos and the other is the one for interbank repos. Both mainly use Treasuries as general collateral. The trading volume of the exchange-traded repo market was less than half of the interbank repo market in 2003. The repo rates prevailing in the exchange traded repo market differ at times from those of the interbank market and are more affected by temporary factors in the equity market, such as new issues. Therefore, we focus on the interbank repo market. Figure 1 shows the total amounts of interbank borrowing and lending without collateral and with collateral (repo). As we can see, repo transactions have surpassed transactions without collateral since 1999 as the most important form of interbank borrowing and lending transactions, totaling more than ten trillion RMB in 2002 and 2003. As the problem of non-performing loans lingers in the Chinese banking industry, the rise of the interbank repo market is no surprise.

Figure 1 here

The interbank repo rate data used in this study are from the www.china-money.com.cn website. The sample includes weekly observations of the one-week rate, Y_t^1, the two-week rate, Y_t^2, the three-week rate, Y_t^3, the one-month (4-week) rate, Y_t^4, the two-month (9-week) rate, Y_t^9, and the three-month (13-week) rate, Y_t^{13}, from July 2, 1999 to June 25, 2004. Although the interbank repo market began in 1997, trading volume before 1999 was too small to warrant study. Table 1 gives annual trading volumes for all the categories by the term of the repo over the period of 1999-2003 and their proportions in the entire interbank repo market.

$\boxed{\text{Table 1 here}}$

The statistics in the table show that, unlike the case of the US repo market in which the overnight repo is most popular, the most popular repo category in China is the one-week repo. The one-week repo has the highest proportion among all the categories. The rate of increase is also the highest for the one-week repo except in 2003. Since 2000, the one-week repo has accounted for more than 60% of the entire interbank repo market. For the categories included in the table, the trading volume decreases with the term of the repo. There are also other terms of repo transactions that are not included in the table. They are of relatively small importance according to the trading volume.

Repo rates are quoted in the Chinese interbank market on the actual/365 basis. In the analysis conducted in this paper, we convert the actual quoted rates to continuously compounded rates. Table 2 reports the descriptive statistics of the repo rates of various terms. Panel A gives the means, standard deviations, and autocorrelations up to the fourth order of the repo rates. Panel B gives the means, standard deviations, and autocorrelations up to the fourth order of the weekly changes of the repo rates.

$\boxed{\text{Table 2 here}}$

As we can see from Panel A of Table 2, the term structure of the repo rates is upward sloped on average. The standard deviation also increases with the term from one week to three months. The pattern in the standard deviations is in contrast with that observed in the US. Longstaff (2001) reports that, while the overnight repo rate has the highest volatility, there is no apparent pattern in the volatilities of the repo rates ranging from one week to three months. The repo rates all have strong and slow decaying autocorrelations.

The numbers in Panel B of Table 2 show that the average weekly changes in the repo rates are very small, less than one basis point for all lengths of the term. The standard deviations are more than 10 basis points, much larger compared with the mean. The rate changes are weakly autocorrelated, compared to the rates themselves.

In order to gain more insight of the analysis for the repo rates that follows, we plot some of the repo rates in Figure 2. All the repo rates are close to each other. To avoid clustering, only the one-week rate, Y_t^1, and the three-month spread, $S_t^{13} = Y_t^{13} - Y_t^1$, are plotted. As seen in the figure, the repo rates drifted slightly down from July 1999 to early 2002. In the second half of 2003, there were two rate spikes. The difference between the rates of different terms is typically small, as evidenced by the three-month term spread in the figure. In late 2003 and early 2004, the one-week rate went through some volatile fluctuations. The interesting observation is that the three-month rate did not fluctuate with the one-week rate on a one-to-one correspondence, resulting in some fluctuation in the term spread. In many cases, the market correctly expected that the fluctuation in the one-week rate was temporary, so the longer-term rates were set much more smoothly over time than were the shorter-term rates.

Figure 2 here

3. Unconditional Term Premiums

The analysis of term premiums is conducted in the backdrop of a discussion of the pure expectation hypothesis. There are many different versions of the pure expectation hypothesis, as explained by Cox et al. (1981). Consequently, there are different definitions of term premiums, i.e., deviations from the pure expectation hypothesis. However, as argued by Campbell (1986), the differences among different versions of the pure expectation hypothesis are of theoretical importance only. The numerical magnitude of such differences are negligible in practice. Fama (1984a, b), Fama and Bliss (1987), and Campbell and Shiller (1991) examine the various notions of term premiums on US Treasuries and reject the pure expectation hypothesis, while Longstaff (2000) studies the repo rates in the US and presents supporting evidence for the pure expectation hypothesis for the short end of the term structure of interest rates. In this paper, we follow Longstaff (2000) and use one of the formulations from Campbell and Shiller (1991).

For the continuously compounded n-week repo rate at t, Y_t^n, one version of the pure expectation hypothesis says that it equals the expected n-week average of one-week rates in the future, $R_t^n = \frac{1}{n}\sum_{j=1}^{n} Y_{t+j-1}^1$. That is,

$$E[Y_t^n - R_t^n] = 0. \tag{1}$$

Note that the one-month, two-month and three-month rates are not exactly 4-week, 9-week and 13-week rates, so the matching here is not perfect. But since some have more days and others have fewer days, the difference should not cause systematic bias, given that they are all annualized rates.

Panel A of Table 3 is for the entire sample period of 1999.07.02—2004.06.25. It reports the sample mean of Y_t^n, R_t^n and $Y_t^n - R_t^n$ with the autocorrelation and conditional heteroskedasticity consistent t-statistics suggested by Newey and West (1987). The lag in the autocorrelation adjustment is chosen to be 30 weeks to account for the high autocorrelations in Y_t^n and the overlapping components in R_t^n. The results show that, on average, the longer repo rates are higher than the corresponding rolling-over short rates, i.e., $Y_t^n - R_t^n$ tends to be positive. As indicated by the t-statistics, the difference in the averages as an estimate of the term premium is significantly positive for all the terms considered in this paper and the term premium increases with the length of the term. The magnitude of the term premiums is small, however. The largest one is about only 21 basis points on the annual basis that occurs for the three-month ($n = 13$) term premium.

Table 3 here

Figure 2 reveals that, over the entire sample period, there is a slight downward trend in the level of repo rates. If the downward trend is totally unexpected, then it is likely to observe a positive realized average value of $Y_t^n - R_t^n$ even when the pure expectation hypothesis holds. To see whether or not this is the case, we look into subperiod results. Panels B and C of Table 3 report the same statistics for two subperiods: 1999.07.02—2002.08.30 and 2002.09.06—2004.06.25. From Figure 2, we can see that the first period is a period in which the level of the repo rates declined overall, while in the second subperiod, the repo rates fluctuated more without following a clear trend. If the positive realized term premiums are entirely due to unexpected realizations, we should anticipate higher realized term premiums in the first subperiod than for those in the entire period and basically no realized term premiums in the second subperiod. The numbers in Panels B and C, however, show the opposite results. The realized term premiums are in fact slightly smaller in the first subperiod than in the second one. This means that the positive realized term premiums found in the entire sample period are not from unexpected changes in the level of repo rates, at least not entirely.

Figure 3 plots the difference $Y_t^n - R_t^n$ for the one-month rate ($n = 4$) and for the three-month rate ($n = 13$). For both series, the difference $Y_t^n - R_t^n$ remains positive for almost all the time. The negative spike of the difference $Y_t^n - R_t^n$ in late 2003 is due to the large temporary fluctuation of the level of the rates. Overall, while the repo rates in early 2004 fall back to the 2002 level, the realized term premiums are positive on average.

Figure 3 here

4. Conditional Term Premiums

The pure expectation hypothesis can also be tested at the conditional level. Stating the hypothesis as

$$E_t[Y_t^n - R_t^n] = 0, \tag{2}$$

where E_t represents the expectation conditioned on time t information, the hypothesis says that the value of $Y_t^n - R_t^n$ should not be predicted by any variable known at t. As a result, one can test the conditional version of the pure expectation hypothesis by regressing $Y_t^n - R_t^n$ on time t variables and testing whether the regression coefficients are all zero. The conditional version of the pure expectation hypothesis is stronger than the unconditional version and, therefore, is easier to reject. Given that the unconditional version of the hypothesis is rejected in the last section for the Chinese interbank repo rates, it seems that testing the conditional version of the hypothesis is beating the dead horse. This view, however, is too scholastic. The purpose of testing a tightly specified economic theory should not be the end of the academic inquiry, but, rather, the beginning point of discussing broader issues surrounding the extreme case represented by the theory. Tests of the conditional version of the pure expectation hypothesis can provide important information about the determinants of conditional term premiums. Indeed, many early tests of the pure expectation hypothesis are of the nature of conditional tests.

In choosing conditioning variables, we look into the modern theory of the term structure of interest rates. There are many term structure models that imply nonzero term premiums. In the most popular one-factor models, such as those by Vasicek (1977) and Cox et al. (1985), the instantaneous rate serves as the only factor. Longstaff and Schwartz (1992) add the conditional volatility of the instantaneous rate as a second factor. More recently, Duffie and Kan (1996) use several key rates as the factors of the entire term structure and as the determinants of term premiums.

To introduce conditioning variables, we estimate a vector autoregressive (VAR) model with the error terms following an exponential generalized autoregressive and conditional heteroskedasticity (EGARCH) model. More specifically, we choose the vector X_t to be

$$X_t = (Y_t^1, S_t^2, S_t^3, S_t^4, S_t^9, S_t^{13})', \tag{3}$$

where $S_t^n = Y_t^n - Y_t^1$ for $n = 2, 3, 4, 9, 13$ is the term spread. The VAR model for X_t is

$$X_t = \mu + \sum_{j=1}^{p} C_j X_{t-j} + \eta_t. \tag{4}$$

The error term, η_t, is assumed to be cross correlated. More specifically, η_t is assumed to be transformed from an uncorrelated vector, e_t, as

$$\eta_t = F e_t = \begin{pmatrix} 1 & 0 & 0 & 0 & 0 & 0 \\ f_{21} & 1 & 0 & 0 & 0 & f_{26} \\ f_{31} & 0 & 1 & 0 & 0 & f_{36} \\ f_{41} & 0 & 0 & 1 & 0 & f_{46} \\ f_{51} & 0 & 0 & 0 & 1 & f_{56} \\ f_{61} & 0 & 0 & 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} e_{1t} \\ e_{2t} \\ e_{3t} \\ e_{4t} \\ e_{5t} \\ e_{6t} \end{pmatrix}. \tag{5}$$

The special feature of matrix F gives η_t a factor structure. In such a structure, e_1 and e_6 are two factors. $e_{1t} = \eta_{1t}$ is the unexpected part of the one-week rate, which is one of the driving forces of the entire term structure of interest rates. As we can see, e_{1t} also affects η_{2t} to η_{6t}. e_{6t} is also a factor in the sense that its value affects other rates except for the one-week rate. It has most influence on η_{6t} if f_{i6} are all small in absolute value. e_{it} for $i = 2, 3, 4$ and 5 is idiosyncratic in the sense that it only affects η_{it}. The conditional variance of e_{it} is denoted by ω_{it} and is assumed to follow a scalar EGARCH(1,1) model,

$$\ln \omega_{it} = \theta_i + \alpha_i e_{i,t-1}/\sqrt{\omega_{i,t-1}} + \beta_i |e_{i,t-1}|/\sqrt{\omega_{i,t-1}} + \gamma_i \ln \omega_{i,t-1}, \qquad i = 1, \cdots, 6. \tag{6}$$

A similar model of this form with a factor structure for the error term is adopted by Bekaert et al. (1997) and Longstaff (2000), among others. Since the number of parameters in the system is too large, the VAR model (4) is estimated first to obtain the realized η_{it}s. The subsystem (5)-(6) is then estimated jointly by the quasi-maximum likelihood approach. To begin the iteration in (6), the expected value or sample average is used for functions of e_{i0} for ω_{i0}. The results of the estimation of (4)-(6) are reported in Table 4. The order of the VAR model is chosen as 5, determined by the likelihood ratio test. The autoregressive part of the VAR model, i.e., μ and C_js, is not reported here because there are too many parameters with little information.

Table 4 here

Panel A of Table 4 reports the estimates of the EGARCH parameters with the t-statistics in parentheses. The autoregressive coefficients, γ_i, are large for all the equations and some of them are close to one, implying a very long half-life of the shocks to the conditional variance. One of the advantages of the EGARCH model is its flexibility in allowing asymmetric responses to positive and negative shocks. Compared to no shock ($e_{i,t-1} = 0$), a positive shock of one standard deviation increases the log conditional variance by $\beta_i + \alpha_i$, while a negative shock of one standard deviation increases the log conditional variance by $\beta_i - \alpha_i$. The β_i estimates are all positive and mostly significant. The α_i estimates are either significantly positive or insignificant. For the equations with significant positive α_i, a positive shock in the term spread causes its conditional variance to increase more than a negative shock of the same magnitude. It should be noted that, although EGARCH(1,1) is very popular in financial econometrics, we do not claim that it is the best model for the Chinese interbank repo rates. We tried several variants and the results do not substantially change the analysis that follows.

Panel B of Table 4 reports the estimates of f_{ij}s with the t-statistics in parentheses. From these estimates, it is seen that the unexpected one-week repo rate, e_{1t}, has a positive influence on the two-week and three-week unexpected repo rates, but a negative influence on the repo rates of maturities more than one month. (Recall that the dependent variables are term spreads, $Y_t^n - Y_t^1$. The negative sign for some of the f_{1j} estimates comes from the $-Y_t^1$ term.) Likewise, e_{6t} has its greatest influence on the unexpected three-month term spread, less on the two-month and one-month unexpected term spreads, and little on the two-week and three-week unexpected term spreads.

To examine the adequacy of the VAR-EGARCH model, we calculate a few statistics of fitted η_{it} and $\tilde{e}_{it} = e_{it}/\sqrt{\omega_{it}}$. The latter is to see the advantage of the EGARCH model. Panel C reports the Bera-Jarque (BJ) statistic test of normality of η_{it} and the Ljung-Box (LB_5) statistics are tests of zero autocorrelation of η_{it} and of η_{it}^2 up to order

5. The BJ statistics show that the normality is strongly rejected for all the equations. The LB_5 statistics for η_{it} show that the fifth-order VAR removes the autocorrelation of the original variables of X_t and the error terms η_{it} are no longer autocorrelated. The LB statistics for e_{it}^2 show that they are autocorrelated and a model with conditional heteroskedasticity is warranted. Panel D reports the Bera-Jarque (BJ) statistic test of normality of $\tilde{e}_{it}$ and the Ljung-Box (LB_5) statistics are tests of zero autocorrelation of $\tilde{e}_{it}$ and of $\tilde{e}_{it}^2$ up to order 5. The BJ statistics in Panel D are much smaller than those in Panel C in most cases. The normality, however, is still strongly rejected. The zero autocorrelation of $\tilde{e}_{it}$ is anticipated given the results in Panel C. The LB_5 statistics for $\tilde{e}_{it}^2$ show that they are no longer autocorrelated and the EGARCH(1,1) model does a good job in removing the autocorrelation that appears in η_{it}^2. Panel E reports the adjusted R^2s for the VAR equations. They indicate that the models fit reasonably.

From the VAR-EGARCH model, we obtain the estimated conditional volatility of the one-week rate, $H_t^1 = \omega_{1,t+1}$, which is the variance of Y_{t+1}^1 conditioned on time t information. We estimate the following regression models:

$$Y_t^n - R_t^n = a_n + b_n Y_t^n + c_n H_t^1 + \varepsilon_t^n, \tag{7}$$

$$Y_t^n - R_t^n = a_n + b_n S_t^n + c_n H_t^1 + \varepsilon_t^n. \tag{8}$$

The choice of Y_t^n follows that of Longstaff (2000) and is motivated by Duffie and Kan's (1996) model. The choice of S_t^n comes from the idea of matching the left-hand side with a rate difference.[1] The choice of H_t^1 is motivated by Longstaff and Schwertz (1992). The regression coefficients are estimated by OLS. The t-ratios of the coefficient estimates are adjusted by the Newey-West scheme using a lag of 30 weeks. The results are reported in Table 5.

$\boxed{\text{Table 5 here}}$

[1]The choice of S_t^n in (8) can also be compared with Fama (1984) in which the left-hand side is the excess holding period return and the right-hand side is the forward spread.

From Table 5, we can see that Y_t^n, S_t^n and H_t^1 all have predictive power for $Y_t^n - R_t^n$. The explanatory power of Y_t^n is small for $n = 2$, but increases substantially for $n > 2$. The explanatory power of S_t^n is quite high for all the n. The result shows that the conditional version of the pure expectation hypothesis does not hold. The conditional term premiums are related to the rates themselves, to term spreads, and to the conditional variance of the one-week rate.

5. Further Analysis

One of the econometric issues in testing the conditional version of the pure expectation hypothesis is the potential bias of the parameter estimates caused by the strong persistence in interest rates. The inference based on asymptotic distributions of the estimators can be erroneous. This can cause over-rejection of the null hypothesis that the conditional term premium is zero. Bekaert et al (1997) give a thorough analysis of the bias issue and present a simulation approach to deal with the problem. Longstaff (2000) adopts Bekaert et al.'s (1997) methodology in his analysis of the US repo market. The bias issue is more important in our case than in Lonstaff's because his evidence is in favor of the pure expectation hypothesis while our results in the previous section are not. The analysis below follows the approach of Bekaert et al. (1997) and Longstaff (2000) in principle.

The simulation is based on the VAR-EGARCH model we build in the last section. In each replication of the simulation, we bootstrap a sample of e_{it}^* for $i = 1, \cdots 6$ and $t = 1, \cdots, T$ from the actual realizations of e_{it}s, where T is the sample size of the actual realizations and an asterisk signifies a simulated value. Using actual observations of $X_0, X_{-1}, \cdots, X_{-4}$ and the estimates of $\mu, C_1, \cdots, C_5, F, \gamma_i, \alpha_i, \beta_i$ for $i = 1, \cdots, 6$ from the last section, we obtain simulated X_{it}^* and ω_{it}^* for $i = 1, \cdots 6$ and $t = 1, \cdots, T$. $Y_t^{1*} = X_{1t}^*$ is the simulated one-week rate. $H_t^{1*} = \omega_{1t}^*$ is the simulated conditional variance of the one-week rate. Simulated values of $(X_{2t}^*, \cdots, X_{6t}^*)$ are discarded. Their role is limited to obtaining simulated Y_t^{1*} and H_t^{1*}, which have the same statistical properties as the actual ones. Instead, we simulate repo rates of longer maturities under the pure expectation hypothesis. At each t, $E_t Y_{t+n}^{1*}$ can be calculated from the VAR model for $n = 1 \cdots 12$. Under the pure expectation hypothesis, $Y_t^{n*} = \frac{1}{n} \sum_{i=0}^{n-1} E_t Y_{t+i}^{1*}$ for $n = 2, 3, 4, 9, 13$. From Y_t^{n*}, we calculate S_t^{n*}. We also calculate the average roll-over return, $R_t^{n*} = \frac{1}{n} \sum_{i=0}^{n-1} Y_{t+i}^{1*}$, for $n = 2, 3, 4, 9, 13$. Finally, regressions (7)-(8) are run for the simulated rates:

$$Y_t^{n*} - R_t^{n*} = a_n^* + b_n^* Y_t^{n*} + c_n^* H_t^{1*} + \varepsilon_t^*, \tag{9}$$

$$Y_t^{n*} - R_t^{n*} = a_n^* + b_n^* S_t^{n*} + c_n^* H_t^{1*} + \varepsilon_t^*, \tag{10}$$

to obtain the estimate of (a_n^*, b_n^*, c_n^*) for $n = 2, 3, 4, 9, 13$.

For the procedure described above, we run 5000 replications. For each $n = 2, 3, 4, 9, 13$ and each equation above, there are 5000 estimates of (a_n^*, b_n^*, c_n^*), which form an empirical distribution of the regression coefficients under the pure expectation hypothesis. Table 6 reports the mean and standard deviation of the empirical distribution of the regression coefficients. For each $n = 2, 3, 4, 9, 13$, the first number is the mean value of a_n^*, b_n^*, or c_n^*. The numbers in parentheses below the means are the standard deviations of the distributions. The p-values of the estimates in Table 5 based on actual observations against the distributions from the simulations are given in the brackets below the standard deviations.

Table 6 here

From Panel A of the table, which gives the results for the equation using the term rate and the conditional volatility as predictors, we can see that all the slope coefficients have positive means for data simulated under the pure expectation hypothesis. In other words, there is some bias to find positive slope coefficients even though the pure expectation hypothesis is correct and the true slope coefficients are zero. The mean coefficients are all smaller than the standard deviation of the distributions. Are the estimated slope coefficients in Table 5 positive because of this bias, then? We can answer this question by examining the p-values of the these estimated coefficients from actual data against the empirical distributions of simulated data. We see that the p-value of b_n, the slope coefficient of Y_t^n from the actual data, is generally small, less than 0.01 in all cases. The p-value of c_n, the slope coefficient of H_t^1, is less than 0.05 for $n = 2$, less than 0.01 for $n = 2, 3$, but is larger for $n = 9$ or 13.

The results in Panel B of Table 6 tell a similar story for the regression using the term spread and the conditional variance as predictors of the term premiums. The mean values of the simulated slope coefficients are small, but not necessarily positive. They are all smaller than the standard deviations in absolute value. The p-values of b_n, the slope coefficient of S_t^n from the actual data, is smaller than 0.0001 in all cases. The p-value of c_n, the slope coefficient of H_t^1, is less than 0.05 for $n = 2$, less than 0.01 for $n = 2, 3$, but is larger for $n = 9$ and 13, very much like the results in Panel A.

Overall, the simulation results indicate that the bias can not be the only reason we find Y_t^n, S_t^n and H_t^1 predictive in Table 5. Y_t^n and S_t^n are indeed part of the determinants of term premiums for all maturities. The conditional variance of the short-term rate, H_t^1, is also part of the determinants of the term premiums, although its role diminishes as the maturity increases.

6. Conclusions

In this paper, we study the behavior of the Chinese interbank repo rates with terms from one week to three months. The repo rates have been the benchmarks for market-determined short-term interest rates in Chinese financial markets. Casual observations reveal that the market has reached a certain level of sophistication. To some degree, the longer-term repo rates reflect the expectation of future fluctuation of the shorter-term repo rates.

We analyze one version of term premiums as deviations of the term rates from the pure expectation hypothesis. The results show that the term premiums are positive and increases with the length of the term. The pure expectation hypothesis is statistically rejected both conditionally and unconditionally. However, the magnitude of the term premiums is economically small. The term premiums are found to be related to the term repo rates themselves, to term spreads, and to the volatility of the one-week repo rate.

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Table 1
Trading volume and market share of repos by term
Panel A of this table presents the trading volume of repos categorized by the term of the transactions. Panel B gives their market share in percentage. The sample period is 1999-2003.

A. Trading volume (billion RMB).

Category	1999	2000	2001	2002	2003
One-week	163.7	1054.0	3092.8	8227.0	7722.4
Two-week	93.4	239.0	470.4	1169.4	1320.4
Three-week	26.7	106.8	189.7	204.1	208.2
One-month	47.2	51.3	88.6	127.6	167.6
Two-month	31.5	65.3	49.7	75.2	86.1
Three-month	26.4	30.8	24.7	41.0	60.4
Entire market	394.9	1578.2	3924.3	9878.9	11306.0

B. Market share (%).

Category	1999	2000	2001	2002	2003
One-week	41.5	66.8	78.8	83.3	68.3
Two-week	23.7	15.1	12.0	11.8	11.7
Three-week	6.8	6.8	4.8	2.1	1.8
One-month	12.0	3.3	2.3	1.3	1.5
Two-month	8.0	4.1	1.3	0.8	0.8
Three-month	6.7	2.0	0.6	0.4	0.5
Sum	98.5	98.0	99.8	99.7	84.6

Table 2
Descriptive statistics of the repo rates
Panel A of this table presents the means, standard deviations, and autocorrelations up to order 4 of the repo rates. Panel B presents the mean, standard deviation, and autocorrelations up to order 4 of the weekly changes in repo rates. The sample consists of weekly observations in the period 1999.07.02–2004.06.25.

A. Rates (%).

Term, n	Mean	St.dev	ρ_1	ρ_2	ρ_3	ρ_4
One-week, Y_t^1	2.3105	0.2654	0.8989	0.7643	0.6794	0.6136
Two-week, Y_t^2	2.3385	0.2829	0.8789	0.7803	0.6664	0.5914
Three-week, Y_t^3	2.3778	0.3016	0.8461	0.7438	0.6686	0.6170
One-month, Y_t^4	2.4368	0.3139	0.8822	0.7825	0.7131	0.6771
Two-month, Y_t^9	2.4764	0.3459	0.8858	0.7682	0.7133	0.7014
Three-month, Y_t^{13}	2.5299	0.3459	0.8867	0.7958	0.7679	0.7294

B. Weekly changes in rates (%).

Term, n	Mean	St.dev	ρ_1	ρ_2	ρ_3	ρ_4
One-week, $Y_t^1 - Y_{t-1}^1$	-0.0010	0.1169	0.1873	-0.2614	-0.1016	0.0762
Two-week, $Y_t^2 - Y_{t-1}^2$	-0.0007	0.1371	-0.0979	0.0711	-0.1732	0.0145
Three-week, $Y_t^3 - Y_{t-1}^3$	-0.0001	0.1651	-0.1885	-0.0737	-0.0828	0.0651
One-month, $Y_t^4 - Y_{t-1}^4$	0.0006	0.1464	-0.0978	-0.1096	-0.1701	0.1024
Two-month, $Y_t^9 - Y_{t-1}^9$	0.0015	0.1554	0.0118	-0.2451	-0.1099	0.0584
Three-month, $Y_t^{13} - Y_{t-1}^{13}$	0.0010	0.1533	-0.0142	-0.3035	-0.0179	0.1349

Table 3
Unconditional term premiums in the repo rates
This table presents the average repo rates of various terms, $\frac{1}{T}\sum_{t=1}^{T} Y_t^n$, for n equal to two weeks, three weeks, one month, two months, and three months. In comparison is the average of corresponding average one-week repo rates, $\frac{1}{T}\sum_{t=1}^{T} R_t^n$. The term premium in column four refers to their difference. The Newey-West t-ratios are used to test the hypothesis that term premium is zero. The sample size is the number of weeks, T, in the sample.

A. sample period for t: 1999.07.02–2004.04.02					
Term, n	Term rates	1-week rate	Term premium	Newey-West	Sample size
	Y_t^n	R_t^n	$Y_t^n - R_t^n$	t-statistics	T
Two-week	2.3368	2.3119	0.0249	3.6060	249
Three-week	2.3705	2.3106	0.0599	3.6907	249
One-month	2.4217	2.3098	0.1120	4.6609	249
Two-month	2.4627	2.3059	0.1568	4.3515	249
Three-month	2.5156	2.3033	0.2123	5.0175	249
B. sample period for t: 1999.07.02–2002.08.30					
Term, n	Term rates	1-week rate	Term premium	Newey-West	Sample size
	Y_t^n	R_t^n	$Y_t^n - R_t^n$	t-statistics	T
Two-week	2.3352	2.3196	0.0156	4.2148	166
Three-week	2.3570	2.3179	0.0391	5.8306	166
One-month	2.3978	2.3166	0.0811	7.4061	166
Two-month	2.4285	2.3112	0.1173	5.5984	166
Three-month	2.4902	2.3069	0.1834	6.2590	166
C. sample period for t: 2002.09.06–2004.04.02					
Term, n	Term rates	1-week rate	Term premium	Newey-West	Sample size
	Y_t^n	R_t^n	$Y_t^n - R_t^n$	t-statistics	T
Two-week	2.3399	2.2965	0.0433	3.2554	83
Three-week	2.3975	2.2960	0.1015	2.9553	83
One-month	2.4697	2.2961	0.1737	3.3341	83
Two-month	2.5311	2.2952	0.2359	2.9200	83
Three-month	2.5662	2.2960	0.2702	2.7086	83

Table 4
A VAR-EGARCH model
This table presents the parameter estimates and certain statistics of the VAR-EGARCH model of the variable $X_t = (Y_t^1, S_t^2, S_t^3, S_t^4, S_t^9, S_t^{13})'$ where $S_t^n = Y_t^n - Y_t^1$,

$$\begin{aligned}
X_t &= \mu + \sum_{j=1}^{5} C_j X_{t-j} + \eta_t, \\
\eta_t &= F e_t = \begin{pmatrix} 1 & 0 & 0 & 0 & 0 & 0 \\ f_{21} & 1 & 0 & 0 & 0 & f_{26} \\ f_{31} & 0 & 1 & 0 & 0 & f_{36} \\ f_{41} & 0 & 0 & 1 & 0 & f_{46} \\ f_{51} & 0 & 0 & 0 & 1 & f_{56} \\ f_{61} & 0 & 0 & 0 & 0 & 1 \end{pmatrix} e_t, \\
\mathrm{Var}_{t-1}[e_t] &= \Omega_t = \mathrm{Diag}(\omega_{1t}, \cdots, \omega_{6t}), \\
\ln \omega_{it} &= \theta_i + \alpha_i e_{i,t-1}/\sqrt{\omega_{i,t-1}} + \beta_i |e_{i,t-1}|/\sqrt{\omega_{i,t-1}} + \gamma_i \ln \omega_{i,t-1}, \qquad i = 1, \cdots, 6.
\end{aligned}$$

Panel A reports the estimates of the EGARCH(1,1) parameters with t-statistics in parentheses. Panel B reports the estimates of f_{ij}s with t-statistics in parentheses. Panel C reports the Bera-Jarque (BJ) statistic test of normality of η_t and the Ljung-Box (LB_5) tests of zero autocorrelation of η_t and of η_t^2 up to order 5, with p-values in brackets. Panel D reports the Bera-Jarque (BJ) statistic test of normality of $\tilde{e}_{it} = e_t/\sqrt{\omega_{it}}$ and the Ljung-Box (LB_5) tests of zero autocorrelation of $\tilde{e}_{it}$ and of $\tilde{e}_{it}^2$ up to order 5, with p-values in brackets. Panel E reports adjusted R^2 for each equation. The sample period is 1999.07.02—2004.06.25.

A. Estimates of EGARCH parameters

i	1	2	3	4	5	6
θ_i	-0.2102	-1.8074	-1.2586	-0.6374	-0.1798	-1.6192
	(-1.9012)	(-3.9964)	(-1.8496)	(-4.0923)	(-1.6052)	(-3.5519)
α_i	0.0252	-0.1326	0.3140	0.3176	-0.0334	0.2550
	(0.7818)	(-1.0967)	(4.5345)	(3.8029)	(-0.5775)	(3.1846)
β_i	0.1780	0.7514	0.2772	0.1002	0.2250	0.3712
	(3.4153)	(5.1307)	(1.5816)	(1.8616)	(2.9575)	(4.0070)
γ_i	0.9809	0.8040	0.8015	0.8926	0.9928	0.7234
	(56.5170)	(13.0249)	(7.6402)	(30.1081)	(58.0776)	(8.0379)

B. Estimates of f_{ij}s

-	f_{21}	f_{31}	f_{41}	f_{51}	f_{61}
-	0.0329	0.1211	-0.0711	-0.4206	-0.3714
-	(1.3139)	(2.1947)	(-1.5929)	(-6.1864)	(-5.7530)
-	f_{26}	f_{36}	f_{46}	f_{56}	-
-	0.0051	0.0796	0.2335	0.3129	-
-	(0.3579)	(1.9952)	(6.0630)	(5.1967)	-

Table 4 (cont'd)

C. Residual analysis of η_{it}						
i	1	2	3	4	5	6
BJ(η_{it})	2121.0127	1360.9458	955.4983	153.3534	1673.8023	176.7698
	[0.0000]	[0.0000]	[0.0000]	[0.0000]	[0.0000]	[0.0000]
$\text{LB}_5(\eta_{it})$	6.4378	2.6417	1.7137	0.6159	4.3590	0.7194
	[0.2659]	[0.7550]	[0.8872]	[0.9873]	[0.4990]	[0.9819]
$\text{LB}_5(\eta_{it}^2)$	7.7935	19.5957	6.1531	11.5281	2.5720	13.7649
	[0.1680]	[0.0015]	[0.2916]	[0.0419]	[0.7656]	[0.0172]
D. Residual analysis of $\tilde{e}_{it} = e_{it}/\sqrt{\omega_{it}}$						
i	1	2	3	4	5	6
BJ($\tilde{e}_{it}$)	998.6105	272.4580	267.5228	280.0241	811.8464	67.6303
	[0.0000]	[0.0000]	[0.0000]	[0.0000]	[0.0000]	[0.0000]
$\text{LB}_5(\tilde{e}_{it})$	11.2752	4.7421	2.8609	1.4676	6.5117	1.6350
	[0.0462]	[0.4482]	[0.7214]	[0.9168]	[0.2596]	[0.8970]
$\text{LB}_5(\tilde{e}_{it}^2)$	0.2350	1.8016	2.8633	0.9538	1.7998	4.0093
	[0.9987]	[0.8759]	[0.7211]	[0.9662]	[0.8761]	[0.5481]
E. Adjusted R^2 for each equation i						
i	1	2	3	4	5	6
R^2	0.8757	0.3930	0.4815	0.6327	0.6684	0.6868

Table 5
Conditional term premiums in the repo rates
This table presents the regression results for term premiums conditioned on either the term rate, Y_t^n, or the term spread, S_t^n, and the conditional volatility, H_t^1, of the one-week repo rate,

$$
\begin{aligned}
Y_t^n - R_t^n &= a_n + b_n Y_t^n + c_n H_t^1 + \varepsilon_t^n, \\
Y_t^n - R_t^n &= a_n + b_n S_t^n + c_n H_t^1 + \varepsilon_t^n.
\end{aligned}
$$

The conditional volatility, H_t^1, is obtained in the VAR-EGARCH model reported in Table 4. Numbers in the parentheses under the parameter estimates are Newey-West t-ratios. R^2s are the adjusted R-squares. The sample is weekly observations from 1999.07.02 to 2004.06.25.

A. $Y_t^n - R_t^n = a_n + b_n Y_t^n + c_n H_t^1 + \varepsilon_t$				
Term n	a_n	b_n	c_n	R^2
Two-week ($n = 2$)	-0.0984	0.0469	0.0199	0.1109
	(-2.6607)	(2.8487)	(3.7913)	
Three-week ($n = 3$)	-0.2129	0.0985	0.0609	0.3127
	(-2.4691)	(2.5779)	(6.3077)	
One-month ($n = 4$)	-0.4529	0.2093	0.0968	0.4979
	(-3.9080)	(4.1236)	(8.5266)	
Two-month ($n = 9$)	-0.8732	0.3941	0.0984	0.6115
	(-3.9637)	(4.4236)	(7.1551)	
Three-month ($n = 13$)	-0.9996	0.4648	0.0874	0.6748
	(-4.7152)	(5.3796)	(4.2926)	
B. $Y_t^n - R_t^n = a_n + b_n S_t^n + c_n H_t^1 + \varepsilon_t$				
Term n	a_n	b_n	c_n	R^2
Two-week ($n = 2$)	-0.0097	0.8908	0.0235	0.5837
	(-2.9953)	(9.2119)	(5.9541)	
Three-week ($n = 3$)	-0.0171	0.7557	0.0552	0.6548
	(-3.1255)	(6.0294)	(5.2567)	
One-month ($n = 4$)	-0.0191	0.7399	0.0851	0.6363
	(-1.8979)	(10.2493)	(5.6038)	
Two-month ($n = 9$)	-0.0164	0.6080	0.1295	0.4757
	(-0.5729)	(4.4608)	(3.6667)	
Three-month ($n = 13$)	0.0084	0.6495	0.1273	0.5063
	(0.2069)	(4.3601)	(2.9003)	

Table 6
Finite sample properties of the conditional term premiums
This table presents simulation results of finite sample properties of the slope coefficients in the regression of term premiums conditioned on the term rate, Y_t^n, the term spread, S_t^n, and the conditional volatility, H_t^1, of the one-week repo rate,

$$\begin{aligned} Y_t^{n*} - R_t^{n*} &= a_n^* + b_n^* Y_t^{n*} + c_n^* H_t^1 + \varepsilon_t^{n*}, \\ Y_t^{n*} - R_t^{n*} &= a_n^* + b_n^* S_t^{n*} + c_n^* H_t^{1*} + \varepsilon_t^{n*}. \end{aligned}$$

The one-week repo rate, Y_t^{1*}, and the conditional variance of the one-week repo rate, H_t^{1*}, are simulated using the model in the equations (4)-(6). The longer-term repo rates, Y_t^{n*}, are calculated under the pure expectation hypothesis using the VAR model (4) reported in Table 4. $R_t^{n*} = \frac{1}{n}\sum_{i=0}^{n-1} Y_{t+i}^{1*}$. The distributions of the OLS estimates of a_n^*, b_n^* and c_n^* are based on 5000 replications. The numbers without parentheses or brackets are the means of the distributions. The numbers in parentheses below the means are the standard deviations of the distributions. The p-values of the estimates in Table 5 against the empirical distributions from the simulations are given in the brackets below the standard deviations.

A. $Y_t^{n*} - R_t^{n*} = a_n^* + b_n^* Y_t^{n*} + c_n^* H_t^{1*} + \varepsilon_t^{n*}$

Term n	a_n^*	b_n^*	c_n^*	R^2
Two-week ($n = 2$)	-0.0136	0.0043	0.0041	0.0047
	(0.0277)	(0.0119)	(0.0077)	(0.0129)
	[0.0064]	[0.0024]	[0.0322]	[0.0002]
Three-week ($n = 3$)	-0.0291	0.0102	0.0083	0.0142
	(0.0558)	(0.0238)	(0.0158)	(0.0224)
	[0.0032]	[0.0012]	[0.0050]	[0.0000]
One-month ($n = 4$)	-0.0487	0.0160	0.0120	0.0201
	(0.0915)	(0.0391)	(0.0242)	(0.0272)
	[0.0004]	[0.0002]	[0.0038]	[0.0000]
Two-month ($n = 9$)	-0.1128	0.0381	0.0249	0.0510
	(0.2349)	(0.1003)	(0.0554)	(0.0549)
	[0.0042]	[0.0012]	[0.0862]	[0.0000]
Three-month ($n = 13$)	-0.1563	0.0536	0.0319	0.0732
	(0.3458)	(0.1476)	(0.0733)	(0.0737)
	[0.0120]	[0.0060]	[0.1772]	[0.0000]

Table 6 (cont'd)

B. $Y_t^{n*} - R_t^{n*} = a_n^* + b_n^* S_t^{n*} + c_n^* H_t^{1*} + \varepsilon_t^{n*}$

Term n	a_n^*	b_n^*	c_n^*	R^2
Two-week ($n = 2$)	-0.0033	-0.0025	0.0042	0.0040
	(0.0056)	(0.0575)	(0.0070)	(0.0124)
	[0.1232]	[0.0000]	[0.0118]	[0.0000]
Three-week ($n = 3$)	-0.0052	-0.0078	0.0087	0.0108
	(0.0109)	(0.0607)	(0.0144)	(0.0201)
	[0.1358]	[0.0000]	[0.0044]	[0.0000]
One-month ($n = 4$)	-0.0107	-0.0143	0.0121	0.0148
	(0.0176)	(0.0780)	(0.0220)	(0.0235)
	[0.2820]	[0.0000]	[0.0034]	[0.0000]
Two-month ($n = 9$)	-0.0218	-0.0352	0.0241	0.0385
	(0.0404)	(0.1044)	(0.0513)	(0.0470)
	[0.5300]	[0.0000]	[0.0356]	[0.0000]
Three-month ($n = 13$)	-0.0273	-0.0504	0.0296	0.0577
	(0.0544)	(0.1214)	(0.0691)	(0.0636)
	[0.7548]	[0.0000]	[0.0738]	[0.0006]

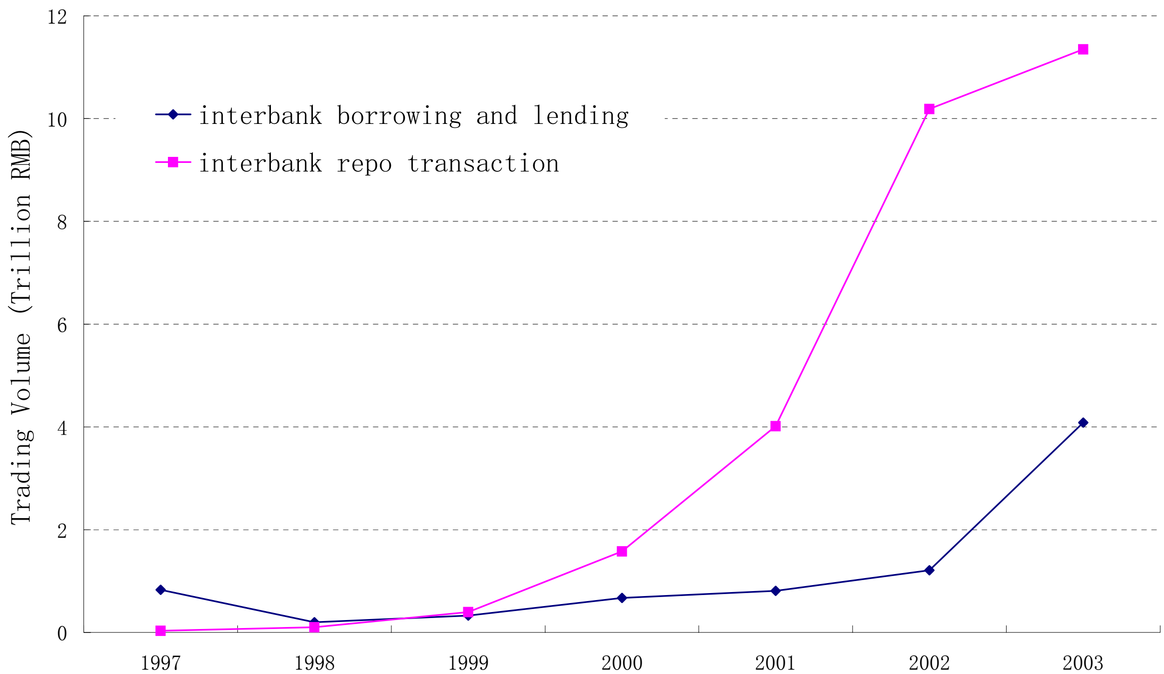

Figure 1. Trading volumes of the interbank borrowing-and-lending market and of the interbank repo market.

This figure shows the annual trading volumes in the two markets from 1999 to 2003.

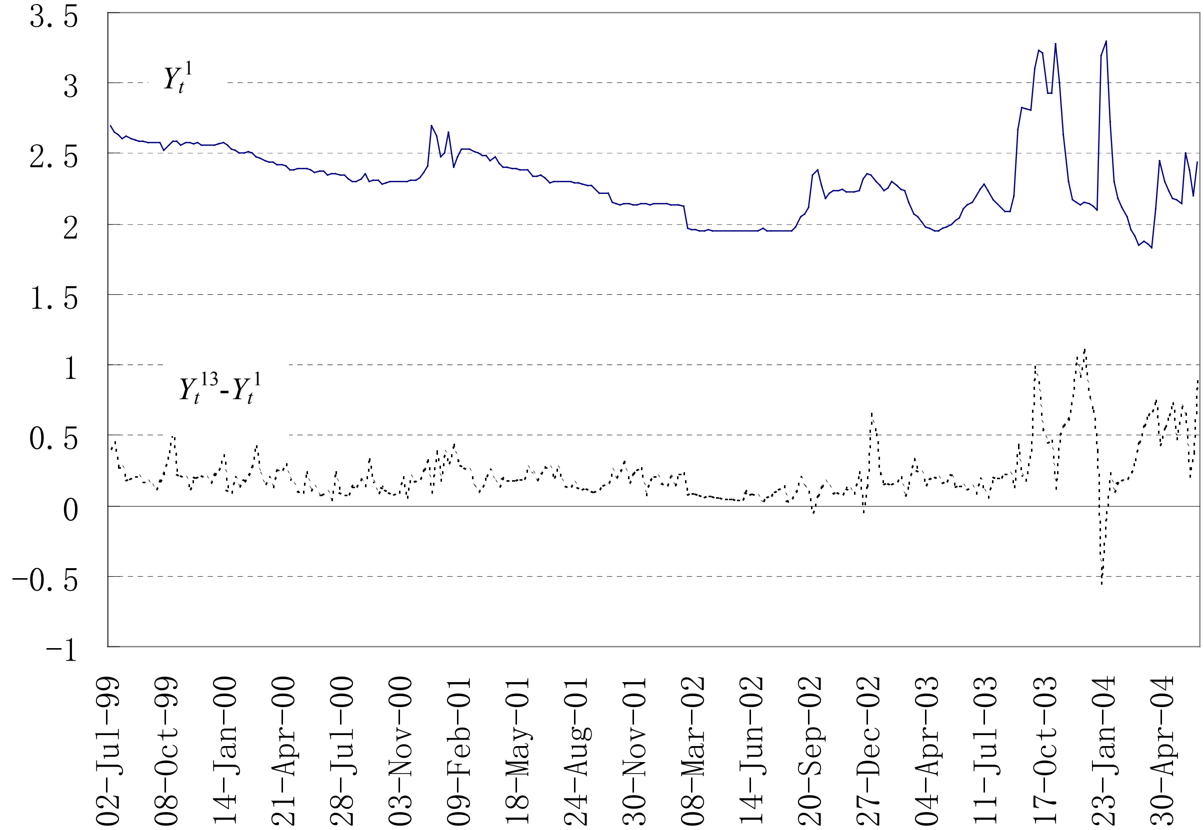

Figure 2. The one-week repo rate and the three-month term spread.

This figure shows the weekly time-series of the one-week repo rate, Y_t^1 (the solid line), and the three-month term spread, $S_t^{13} = Y_t^{13} - Y_t^1$ (the dashed line), from 1999.07.02 to 2004.06.25.

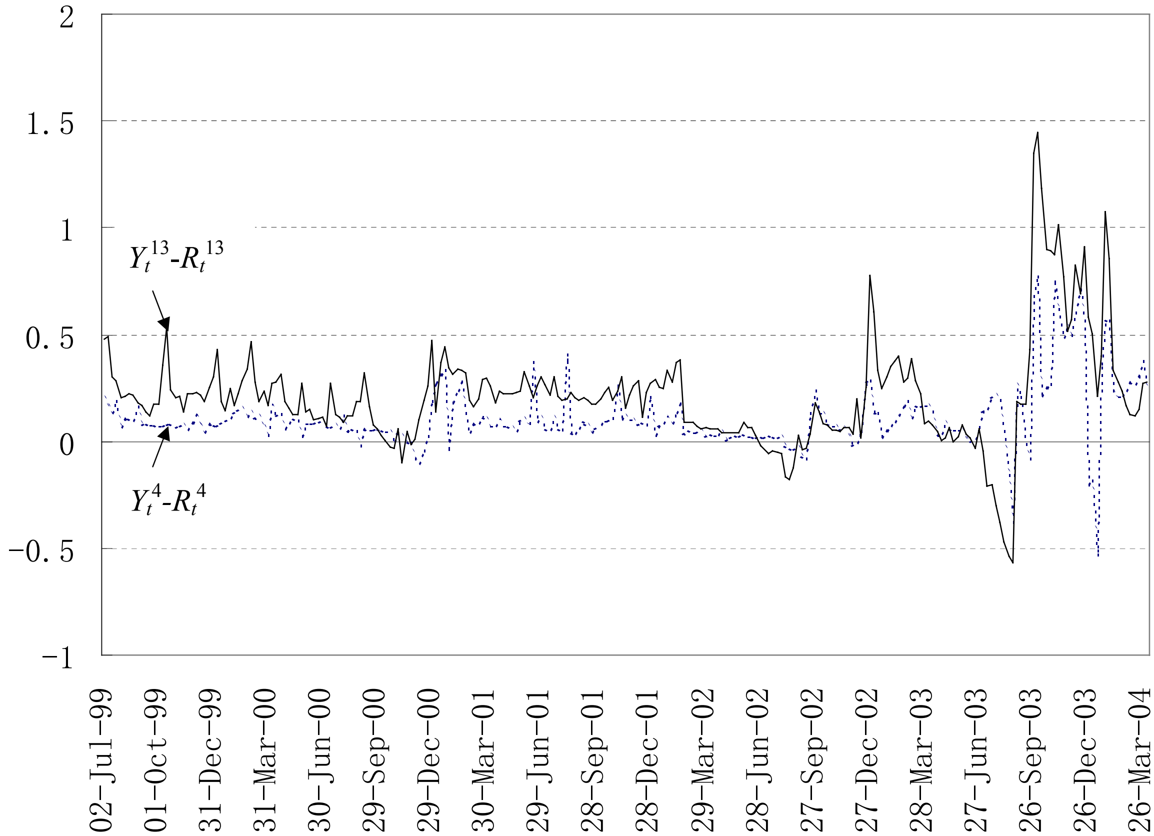

Figure 3. Longer-term repo rates in excess of rolling-over one-week rates. This figure shows the weekly time-series of the one-month repo rate in excess of the rolling-over one-week repo rates, $Y_t^4 - R_t^4$ (dashed line), and the three-month repo rate in excess of the rolling-over one-week repo rates, $Y_t^{13} - R_t^{13}$ (the solid line), from 1999.07.02 to 2004.04.02.