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Pacific-Basin Finance Journal 11 (2003) 593–609
PACIFIC-BASIN FINANCE JOURNAL
www.elsevier.com/locate/econbase

Interpreting the term structure of interbank rates in Hong Kong

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Received 12 November 2001; accepted 19 December 2002

Abstract

This paper studies the term structure of 1-, 3-, 6-, 9- and 12-month interbank rates in Hong Kong using monthly data spanning 1992:01–2000:03. We find that term spreads contain no information about future short-term rates. The Expectations Hypothesis (EH), which states that long-term rates depend on expected future short-term rates plus a constant term premium, is also soundly rejected by the data. However, we are unable to reject a modified version of the EH that incorporates time-varying term premia. Moreover, the hypothesis that the parameters are constant before and after the onset of the Asian crisis in 1997:06 is not rejected.

JEL classification: E43
Keywords: Term structure of interest rates; Expectations Hypothesis; Hong Kong

1. Introduction

Central banks typically pay considerable attention to the term structure of interest rates in their day-to-day monitoring of financial markets.[1] This is so for several reasons. First, there is empirical work using data from a range of economies indicating that the slope of the term structure contains information about economic activity in the near term. Estrella and Hardouvelis (1991) demonstrate that a downward-sloping term structure appears to be

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[1] For a review of the literature on the term structure of interest rates, see Shiller (1990).

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doi:10.1016/S0927-538X(03)00014-3

associated with an increased likelihood of low future economic growth or a recession.[2] Second, the slope of the term structure may be informative about the future path of inflation (see Mishkin, 1990a,b, 1991).[3] Thus, an upward-sloping term structure is commonly interpreted as suggesting that inflation is expected to rise. Third, and arguably of more direct importance to monetary policy makers, the term structure may contain information about market participants' expectations regarding the future path of short-term interest rates. Since monetary policy in most countries, although not in Hong Kong, is conducted by the central bank actively influencing very short-term interest rates, the term structure can thus be seen as capturing market participants' expectations of the future stance of monetary policy. However, while economic theory suggests that longer-term interest rates are influenced by expectations of future short-term interest rates, other factors also play a role in determining interest rates. For instance, shifts in liquidity or market participants' assessment of the riskiness of holding deposits or securities of different maturities may affect the term structure. If the importance of such factors changes over time, the induced variation in liquidity and risk premia can render the interpretation of the term structure more difficult.

In this paper, we study monthly data spanning 1992:01–2001:02 on the term structure of interbank rates in Hong Kong. The purpose of the analysis is to assess to what extent the term structure contains information about the future path of short-term interest rates. In Section 2, we characterise the evolution over time of short-term interest rates and term spreads. As a prelude to the econometric analysis, we test for unit roots in the different interest rates and term spreads.

In Section 3, we test whether the data obey the Expectations Hypothesis (EH) of the term structure, that is, whether we can reject the hypothesis that term premia are constant over time and longer-term interest rates are determined solely by the expected future path of short-term rates. We show that the data reject soundly this hypothesis and conclude that time-varying term premia are present. Indeed, we find that there appears to be no information about the future path of interest rates embedded in the term spreads. In light of this, we incorporate time-varying term premia in the econometric analysis. For simplicity, we model these as given by the conditional variance of innovations to the 1-month interest rate, which we estimate using an ARCH model. Our proxy for the term premia is highly significant in the different regressions, and we find strong evidence that interest rate spreads are informative for the future path of short-term rates. Nevertheless, the EH is rejected by the data. However, since the ARCH model yields at best a noisy estimate of the true volatility, these regressions are subject to errors-in-variables bias of unknown magnitude. We therefore re-estimate the equations, using a moving average of lagged squared changes of the 1-month rate to instrument volatility, as suggested by Pagan and Ullah (1988). In these final estimates, the risk premia remain highly significant. More importantly, we can typically reject the hypothesis that the slope parameters are zero, but are unable to reject the hypothesis that they are unity as suggested by theory. Finally, the hypothesis that the parameters did not change after the onset of the Asian crisis in 1997:06 is not rejected. The absence of signs of a

[2] See also Bernanke (1990), Harvey (1988, 1991), and Kamara (1997). Bernard and Gerlach (1998) provide international evidence.

[3] See also Jorion and Mishkin (1991) and Gerlach (1997).

shift in the parameters despite the fact that interest rate volatility rose sharply suggests that the model explains the determination of the term structure of short-term interest rates in Hong Kong quite well.

Finally, Section 4 contains our conclusions. Overall, the results suggest that the term structure of interbank rates in Hong Kong can be thought of as being driven by expectations of future short-term rates and a term premium that is related to the volatility of the 1-month rate.

2. Preliminaries

In this section, we review the data and present some simple stylised facts with respect to their behaviour. The data used are from the Hong Kong Monetary Authority's website (www.hkma.gov.hk) and represent interbank offered rates with 1, 3, 6, 9 and 12 months maturity. The data are monthly and span the period 1992:01 to 2001:02.

2.1. Implications of the currency board

The Hong Kong dollar (HKD) has been tied to the US dollar (USD) through a currency board arrangement ("the Link") since October 1983.[4] As a consequence, HKD interest rates are closely, but not perfectly, related to those in the United States. Two factors can cause temporary deviations between the two yields. First, expectations of a change in the currency board will lead to a divergence of yields. For instance, Gerlach (2002) shows that HKD yields rose sharply relative to USD rates during the second half of 1998 as the HKD was exposed to heavy speculative pressures. Second, episodes of limited liquidity in HKD markets can render yields uninformative. The HKD market has developed rapidly in the 1990s, but liquidity is still less than infinite as is required for the two yield curves to be perfectly arbitraged.

The operation of the currency board implies that the 1-month rate that we view as driving the evolution over time of the term structure of interest rates in Hong Kong is determined largely by forces external to the domestic economy. It should be noted from the outset that this fact does not have any implications for tests of the EH since, as we discuss further below, it merely states that there is a relationship between interest rates of different maturities, whatever the factors are that determine them. However, while the currency board arrangement has no direct implications for the behaviour of the yield curve, the fact that HKD and USD yield curves are strongly related suggests that tests of the EH would fare about as well in HKD as in USD data. From this perspective it is notable that Hardouvelis (1994), who looks at longer-term bonds yields in the G-7 countries, finds that the EH is most easily rejected in USD data, and argues that this is "puzzling." Further evidence that US yield curve is explained poorly by the EH is provided by Gerlach and Smets (1997). They study the returns on 1-, 3-, 6- and 12-month euro-dollar deposits for 17 countries and find that the

[4] Gerlach (2002) contains a discussion of the currency board arrangement and the day-to-day relationship between HKD and USD very short-term interest rates.

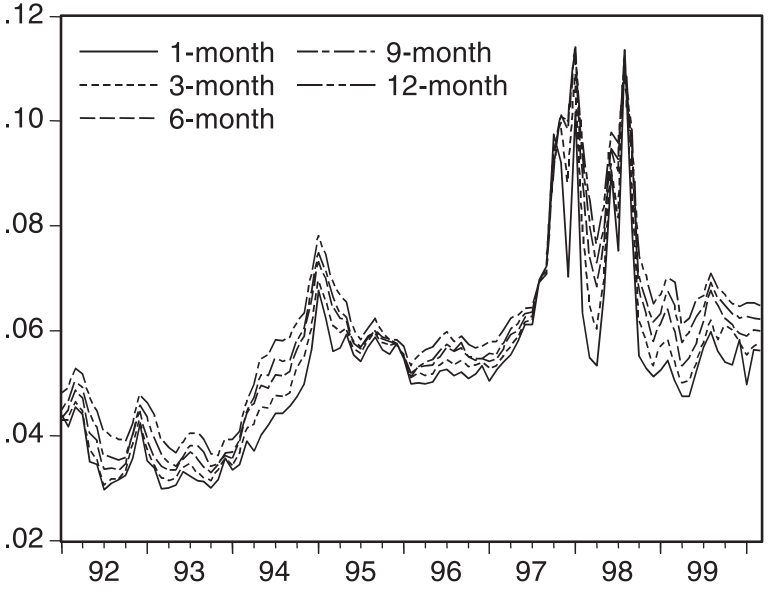

Fig. 1. Interbank rates.

EH is most easily rejected in the Austrian and US data.[5] Overall, these results suggest that the HKD data is also unlikely to be easily reconcilable with the EH.

2.2. The behaviour of HKD interest rates

Fig. 1 shows that the term structure of interbank rates in Hong Kong experienced several episodes of turbulence in the period under study. Particularly noteworthy are the sharp increases in the middle of 1997 following the floating of the Thai baht and in the third quarter of 1998 following the Russian debt moratorium and a massive speculative attack against the Hong Kong currency board. Less pronounced interest rate increases took place in late 1992 at the time of speculative attacks against a number of European currencies pegged to the ECU and in late 1994.

Fig. 1 also indicates that movements in interest rates tend to be associated with changes in term spreads, that is, spreads between interest rates of different maturity. To demonstrate this more clearly, Fig. 2 plots the spreads of 3-, 6-, 9- and 12-month interest rates against the 1-month rate. Two aspects of the figure are of particular interest. First, longer interest rates tend to exceed short-term rates, that is, the term structure is generally upward sloping. Second, movements in the 3–1 month spread are reflected in movements in the longer-term spreads, with the extent of the reactions positively related to the maturity of the long rate used to construct the spread. Thus,

[5] Sutton (2000) shows that the evidence against the EH in US data comes from the period of non-borrowed reserve targeting in 1979–1982.

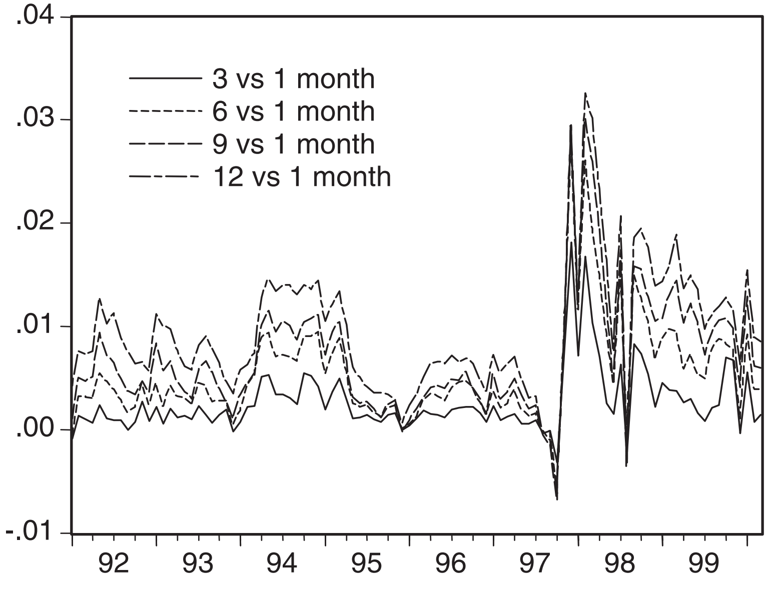

Fig. 2. Term spreads.

the spread between 12- and 1-month rates is more variable than the spread between 9- and 1-month rates, which in turn is more variable than the spread between 6- and 1-month rates.

Table 1 shows the sample means and standard deviations of the term spreads (defined as the longer interest rate minus the 1-month rate). Since we test for a structural break at the onset of the Asian financial crisis in the summer of 1997 in the econometric analysis below, we present results for the full sample, 1992:01–2000:03,

Table 1
Means and standard deviations for interest rate spreads

	Full Sample 1992:01–2000:03			
Spread (× 100)	3–1 month	6–1 month	9–1 month	12–1 month
Mean (× 100)	0.262	0.540	0.698	0.947
Standard deviation (× 100)	0.312	0.500	0.591	0.636
	First subsample 1992:01–1997:05			
Mean (× 100)	0.180	0.387	0.516	0.773
Standard deviation (× 100)	0.136	0.238	0.293	0.349
	Second subsample 1997:06–2000:03			
Mean (× 100)	0.418	0.832	1.046	1.280
Standard deviation (× 100)	0.463	0.707	0.826	0.889

Table 2
Phillips–Perron tests for unit roots

	1992:01–2000:03
	Test statistic
Interest rates	
1-month	−4.06**
3-month	−3.16*
6-month	−2.83
9-month	−2.66
12-month	−2.60
Term spreads	
3–1 month	−6.22**
6–1 month	−5.68**
9–1 month	−5.55**
12–1 month	−5.34**

A constant and a time trend were included in the tests and three lags were used.

* Significance at the 5% level.

** Significance at the 1% level.

and for the subsamples 1992:01–1997:05 and 1997:06–2000:03. The term structure is on average upward sloping in both subsamples, but more sharply so in the sample covering the Asian financial crisis. Furthermore, the volatility is also greater in the second sample. The fact that the term spreads are positively related to the maturity of the long rate used to compute them suggests that term premia are present. However, and as we discuss below, that does not imply that the EH, defined formally in Section 3, is inconsistent with the data.

Next we perform Phillips–Perron tests on the interest rates and the term spreads. This is an important preliminary step to the econometric analysis that follows below since any use of non-stationary data would have implications for the ways in which inference should be conducted. The first column of Table 2 indicates that we can reject the hypothesis of a unit root for the 1- and 3-month rate and for the term spreads. However, for the level of the 6-, 9- and 12-month rates we are unable to reject the hypothesis of a unit root.[6] Since the econometric analysis below is conducted on changes of the 1-month rate and the term spreads, which all are stationary in levels, it follows that hypothesis testing can be carried out using standard procedures.

3. Testing the Expectations Hypothesis

So far, we have merely described the time series behaviour of interbank rates in Hong Kong. In this section, we attempt to interpret the data, using the EH of the term structure as our point of reference. The critical implication of the EH is the testable

[6] The rejection of the unit root hypothesis for 1-month rates is important in light of the finding of Bekaert et al. (1997) that strong persistence of the short rate can lead to a bias in the tests we run below.

assumption that the expected return from investing in alternative segments of the term structure is the same, except possibly for a time-invariant term premium.[7] More formally, let $R_t^{(N)}$ denote the N-period spot rate at time t. The EH can then be written as

$$\left(1+R_t^{(N)}\right)^N = \left(1+E_tR_t^{(1)}\right) \times \left(1+E_tR_{t+1}^{(1)}\right) \times \left(1+E_tR_{t+2}^{(1)}\right) \times \cdots \times \left(1+E_tR_{t+N-1}^{(1)}\right) \times \Theta^{(N)}, \tag{1}$$

where $\Theta^{(N)}$ denotes the term premium. Letting lower case letters denote compounded rates (so that $r_t^{(N)}) = \ln(1+R_t^{(N)})$ and defining $\theta^{(N)} = \ln\Theta^{(N)}$, we have that

$$r_t^{(N)} = \theta^{(N)} + \frac{1}{N}\sum_{i=0}^{N-1} r_{t+i}^{(1)}. \tag{2}$$

To proceed, we subtract $r_t^{(1)}$ from both sides and rearrange

$$\frac{1}{N}\sum_{i=0}^{N-1}\left(E_tr_{t+i}^{(1)} - r_t^{(1)}\right) = -\theta^{(N)} + r_t^{(N)} - r_t^{(1)}. \tag{3}$$

To interpret Eq. (3), consider the simplest case, in which $N=2$

$$\frac{1}{2}\left(E_tr_{t+1}^{(1)} - r_t^{(1)}\right) = -\theta^{(2)} + r_t^{(2)} - r_t^{(1)}, \tag{4}$$

and suppose that the term premium is zero. Eq. (4) then states that there is a linear relationship between the expected change in the 1-month rate, $E_tr_{t+1}^{(1)} - r_t^{(1)}$, and the spread between the 2- and 1-month rates, $r_t^{(2)} - r_t^{(1)}$. Thus, if short-term rates are expected to rise (fall), the term structure will be upward- (downward-) sloping.

Before proceeding, note that Eq. (4) explains why the slope of the term structure is useful for predicting future inflation and output. To understand this, suppose that central banks raise interest rates in response to inflation and that this in turn reduces economic activity. If market participants come to believe that inflation is about to rise, they will expect the central bank to increase short-term rates in the future. According to Eq. (4), this implies that longer-term rates start to rise already now. If, on average, market participants are correct in their assessments of economic developments, we will see an upward-sloping term structure associated with higher future short-term rates, higher future inflation, and lower future economic growth.

One problem with testing the EH is that Eq. (4) contains expectations of future short-term rates and can therefore not be estimated directly. To overcome this problem, note that by definition $r_{t+j}^{(1)} \equiv E_tr_{t+j}^{(1)} + \varepsilon_{t+j}^{(1)}$, which states that the actual 1-month rate is equal to its expected value plus an expectation error (defined by $\varepsilon_{t+j}^{(1)} \equiv r_{t+j}^{(1)} - E_tr_{t+j}^{(1)}$). If we are willing

[7] The term premium may, however, differ between maturities.

to make the standard assumption that the expectation errors are serially uncorrelated, we can use this result to rewrite Eq. (4) as

$$\frac{1}{N}\sum_{i=0}^{N-1}\left(r_{t+i}^{(1)}-r_{t}^{(1)}\right)=-\theta^{(N)}+r_{t}^{(N)}-r_{t}^{(1)}+\frac{1}{N}\sum_{i=0}^{N-1}\varepsilon_{t+i}^{(1)}. \tag{5}$$

Consider next the equation

$$\frac{1}{N}\sum_{i=0}^{N-1}\left(r_{t+i}^{(1)}-r_{t}^{(1)}\right)=\alpha^{(N)}+\beta^{(N)}\left(r_{t}^{(N)}-r_{t}^{(1)}\right)+v_{t}^{(N)}. \tag{6}$$

The EH then holds that $\alpha^{(N)}=-\theta^{(N)}$ and $\beta^{(N)}=1$. Thus, by estimating Eq. (6) and testing whether the slope parameter is unity we can test the EH.[8] Doing so is straightforward, with the minor complication that the error term, $v_t^{(N)}=\frac{1}{N}\sum_{i=0}^{N-1}\varepsilon_{t+i}^{(1)}$, obeys a moving-average structure of order $N-1$.

3.1. Preliminary estimates

We first estimate Eq. (6) using GMM on monthly data spanning 1992:01–2000:03. Standard errors are computed using the approach suggested by Newey and West (1987), which allows the errors to be heteroscedastic and to obey an MA($N-1$) structure. The results in Table 3 are discouraging for the EH in that the slope parameters are close to zero, ranging from 0.16 to 0.37. More importantly, they are significantly different from unity, but, except for the case of $N=9$, not significantly different from zero. Moreover, the constants are all negative, but not significant. Finally, the adjusted R-squared is extremely low in all regressions. Of course, this is merely a reflection of the fact that the turbulence the term structure underwent in the latter part of the sample was unexpected and led to very large expectation errors.

3.2. Time-varying term premia

The fact that the slope parameters are estimated to be between zero and unity suggests that allowing for time-variation in the term premium may be critical for understanding the movements of the term structure of interbank rates in Hong Kong. To see this, rewrite Eq. (3) as

$$E\Delta r_t^{(N)}=-\theta_t^{(N)}+\mathrm{TS}_t^{(N)}, \tag{3$'$}$$

[8] It should be noted that this is merely one of several ways in which the EH hypothesis can be tested. The reason for concentrating on this aspect is that it focuses on the issue of how much information term spreads contain about market participants' expectations of future short rates, which arguably is the most interesting question from the perspective of central banks.

Table 3
GMM estimates of $\frac{1}{N}\sum_{i=0}^{N-1}\left(r_{t+i}^{(1)}-r_t^{(1)}\right)=\alpha^{(N)}+\beta^{(N)}\left(r_t^{(N)}-r_t^{(1)}\right)+v_t^{(N)}$

Maturity of long rate (N)	1992:01–2000:03		
	$\alpha^{(N)}$	$\beta^{(N)}$	$\bar{R}^2$
3 months	− 0.000	0.155	− 0.00
	(0.001)	(0.273)	
	[0.810]	[0.570]	
6 months	− 0.000	0.177	− 0.00
	(0.002)	(0.225)	
	[0.774]	[0.432]	
9 months	− 0.002	0.371	0.04
	(0.003)	(0.157)	
	[0.478]	[0.021]	
12 months	− 0.001	0.242	0.01
	(0.003)	(0.161)	
	[0.730]	[0.136]	

Newey–West standard errors in parentheses; p-values in brackets.

where $E\Delta r_t^{(N)}\equiv\frac{1}{N}\sum_{i=0}^{N-1}\left(E_t r_{t+i}^{(1)}-r_t^{(1)}\right)$ and $\mathrm{TS}_t^{(N)}\equiv r_t^{(N)}-r_t^{(N)}$. Mankiw and Miron (1986) show that in the presence of a time-varying term premium, the estimated value of $\beta^{(N)}$ in Eq. (6) is given by [9]

$$\beta^{(N)}=\frac{\mathrm{Var}(E\Delta r^{(N)})+\mathrm{Cov}(E\Delta r^{(N)},\theta^{(N)})}{\mathrm{Var}(E\Delta r^{(N)})+\mathrm{Var}(\theta^{(N)})+2\mathrm{Cov}(E\Delta r^{(N)},\theta^{(N)})}. \tag{7}$$

Eq. (7) indicates that if the variance of the term premium is zero (implying that the covariance of the term premium and the expected change in the interest rate is zero), the slope parameter is indeed unity. However, if there is a term premium, the slope parameter can be negative or larger than unity. Only if the covariance is not too large, $\beta^{(N)}$ will range between zero and unity.

Next we develop an empirical measure of the term premium. While economic theory suggests that term premia depend on the covariance of the return of the asset in question with the market, we assume for simplicity that the term premium is proportional to the logarithm of the variance of the innovation to the 1-month rate, σ_t^2.[10] We can then estimate an augmented version of Eq. (6)

$$\frac{1}{N}\sum_{i=0}^{N-1}\left(r_{t+i}^{(1)}-r_t^{(1)}\right)=\alpha^{(N)}+\beta^{(N)}\left(r_t^{(N)}-r_t^{(1)}\right)+\delta^{(N)}\log\left(\sigma_t^2\right)+v_t^{(N)}, \tag{8}$$

where $\delta^{(N)}$ captures the impact of the term premium. We first estimate Eq. (8) directly using GMM. One problem in doing so is that since σ_t^2 is not observed, we are required to

[9] See also chapter 10 in Campbell et al. (1997).

[10] See chapter 19 in Cuthbertson (1996). In the working paper version of this paper, it was assumed the term premium was proportional to the variance of changes in interest rates. However, preliminary estimates indicated that using the logarithm of the variance improved the fit sharply. Since theory is silent on the functional form, it seemed preferable to adopt this specification.

Table 4
Value of log-likelihood function for GARCH(i,j) models 1992:01 – 2000:03

	ARCH terms (i)	
	1	2
GARCH terms (j)		
0	406.03	433.69
	[0.000]	[0.720]
1	428.01	433.97
	[0.002]	[0.752]
2	429.71	434.02
	[0.003]	

p-values for test of restrictions against GARCH(2,2) model in brackets.

use an estimate thereof, $\hat{\sigma}_t^2$. Since any such estimate is imperfect, this procedure introduces an errors-in-variables problem that potentially could lead to biased estimates of $\delta^{(N)}$. In the econometric analysis reported on below, we initially disregard this problem. Next, we follow the suggestion of Pagan and Ullah (1988) and use instrumental variables to overcome the measurement error in volatility. Before turning to the results, we first estimate the volatility of the 1-month rate.

3.3. GARCH models

The starting point for the empirical analysis of the term premium is a low-order GARCH model for the monthly innovation in the 1-month rate. Since GARCH(1,1) models tend to capture the volatility of many financial time series, we purposely overfit the data by estimating a GARCH(2,2) model.[11] More precisely, we estimate[12]

$$\Delta r_t^{(1)} = \gamma_0 + \gamma_1 r_{t-1}^{(1)} + \varepsilon_t, \tag{9}$$

where $\varepsilon_t \sim N(0, \sigma_t^2)$ and where

$$\sigma_t^2 = \mu_0 + \sum_{i=1}^{2} \tilde{\mu}_i \sigma_{t-i}^2 + \sum_{j=1}^{2} \mu_j \varepsilon_{t-j}^2. \tag{10}$$

Next we estimate restricted GARCH(i,j) versions of this model and calculate the p-values associated with likelihood ratio tests for the restrictions. The results in Table 4 indicate that a GARCH(0,2)—or, equivalently, an ARCH(2)—model fit the data well, and we therefore use the implied series of $\hat{\sigma}_t^2$ as a measure of the risk premium.[13] Table 5

[11] Allowing for higher-order models leads to occasional convergence problems, suggesting that the models are overfitted.

[12] The model was estimated using the maximum likelihood routine, which assumes conditional normality, implemented in EViews 4.1.

[13] The referee suggested that the square-root augmented version of the (G)ARCH model be explored since Bekaert et al. (1997) provide some evidence in favour of this model using US data (see also Gray, 1996). We nested the square-root model with the ARCH model by introducing the lagged level of the 1-month rate as a determinant of volatility in Eq. (10). However, the lagged interest rate was always insignificant while the two past squared errors remained highly significant. The data thus strongly prefer the ARCH(2) specification.

Table 5
Estimates of

$$\Delta r_t^{(1)} = \gamma_0 + \gamma_1 r_t^{(1)} + \varepsilon_t$$

$$\sigma_t^2 = \mu_0 + \sum_{j=1}^{2} \phi_j \varepsilon_{t-j}^2$$

1992:01–2001:02	
γ_0	0.002
	(0.001)
	[0.070]
γ_1	− 0.043
	(0.025)
	[0.085]
μ_0	2.39e − 06
	(1.46e − 06)
	[0.102]
ϕ_1	0.705
	(0.239)
	[0.003]
ϕ_2	0.853
	(0.190)
	[0.000]
$\bar{R}^2$	0.002
DW	2.55

Newey–West standard errors in parentheses; p-values in brackets.

provides the parameter estimates of this model for the full sample, 1992:01–2001:03. Before proceeding, it is of interest to consider the estimated values for $\hat{\sigma}_t^2$ implied by this model. Fig. 3 indicates that the volatility of the 1-month rate was generally low during the sample, but rose dramatically following the onset of the Asian financial crisis in 1997 and during the latter half of 1998 when the Hong Kong currency board was subject to a massive speculative attack. This suggests that the term premium largely reflects the need for an additional return on Hong Kong dollar assets to induce market participants to hold them in periods of uncertainty regarding the peg.

3.4. Estimates incorporating risk premia

Having obtained an estimate of the variance of shocks to the 1-month rate, $\hat{\sigma}_t^2$, we estimate

$$\frac{1}{N}\sum_{i=0}^{N-1}\left(r_{t+i}^{(1)} - r_t^{(1)}\right) = \alpha^{(N)} + \beta^{(N)}\left(r_t^{(N)} - r_t^{(1)}\right) + \delta^{(N)}\log\left(\hat{\sigma}_t^2\right) + v_t^{(N)} \tag{11}$$

using GMM. The results in Table 6 are more supportive of the EH than those in Table 3. In particular, the estimated β parameters are numerically larger than before. While they are insignificantly different from zero and unity for $N=3$, they are significantly different from zero but insignificantly different from unity when $N=6, 9, 12$. Moreover, while the parameter attached to the volatility of the 1-month rate, which we take as a measure of

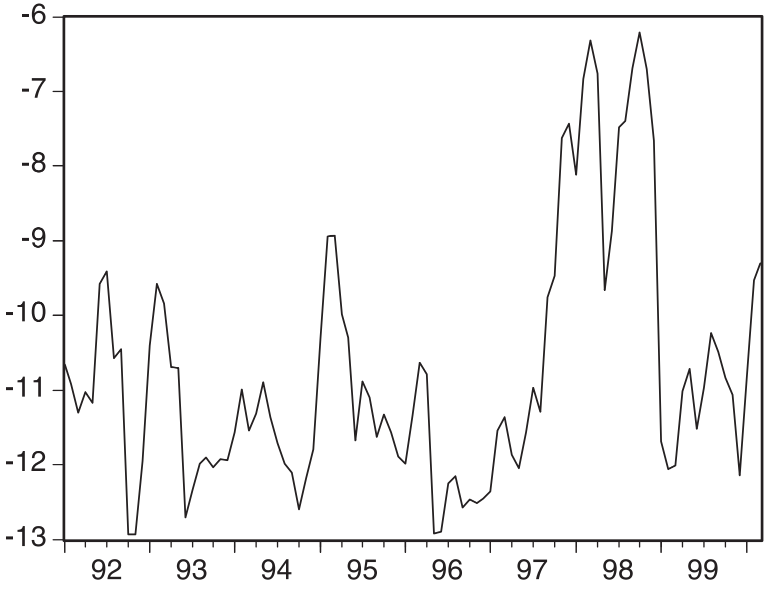

Fig. 3. Estimated volatility of changes in 1-month rates (from ARCH(2) model, in logs).

the term premium, is insignificant for the 3-month rate, it is highly significant in the other cases. Note also that the values of $\bar{R}^2$ are much higher than in Table 3.

While these estimates are encouraging, it should be noted that since we use an estimate of the variance of shocks to the 1-month rate, $\hat{\sigma}_t^2$, rather than the true (unobserved) variance, σ_t^2, the results in Table 6 are subject to generated-regressors bias of unknown importance. Pagan and Ullah (1988) show that this can lead to severely biased parameter estimates and propose an instrumental variables strategy to overcome this problem.[14] We therefore re-estimate these equations following the strategy suggested by Pagan and Ullah (1988, pp. 101–103). Thus, we replace (the logarithm of) σ_t^2 as a regressor with (the logarithm of) the square of the fitted error in the ARCH(2) model for $\Delta r_t^{(1)}$, $\hat{\varepsilon}_t^2$. That is, we estimate

$$\frac{1}{N}\sum_{i=0}^{N-1}\left(r_{t+i}^{(1)} - r_t^{(1)}\right) = \alpha^{(N)} + \beta^{(N)}\left(r_t^{(N)} - r_t^{(1)}\right) + \delta^{(N)}\log\left(\hat{\varepsilon}_t^2\right) + v_t^{(N)}. \qquad (11')$$

Since $\hat{\varepsilon}_t{}^2$ is a noisy measure of the true, unobserved volatility of the 1-month rate, Pagan and Ullah (1988) suggest instrumenting it with the estimated conditional variance, $\hat{\sigma}_t^2$, arising from the ARCH model for $\Delta r_t{}^{(1)}$.

14 As pointed out by the referee, it is possible to estimate the volatility model and the term structure equation jointly. However, while the latter has a structural interpretation, the former does not. In the absence of a structural model, there is a risk that the volatility equation is subject to specification error. Since we are merely exploring the hypothesis that the term premium is correlated with the volatility of the 1-month rate (rather than proposing a formal model for it), we estimate the equations separately to reduce the implications of any specification error.

Table 6
GMM estimates of $\frac{1}{N}\sum_{i=0}^{N-1}\left(r_{t+i}^{(1)}-r_t^{(1)}\right)=\alpha^{(N)}+\beta^{(N)}\left(r_t^{(N)}-r_t^{(1)}\right)+\delta^{(N)}\log\left(\hat{\sigma}_t^2\right)+v_t^{(N)}$

Maturity of long rate (N)	1992:01–2000:03			
	$\alpha^{(N)}$	$\beta^{(N)}$	$\delta^{(N)}\times 100$	$\bar{R}^2$
3 months	− 0.012	0.437	− 0.103	0.042
	(0.010)	(0.338)	(0.079)	
	[0.213]	[0.199]	[0.199]	
		[0.099]		
6 months	− 0.027	0.613	− 0.232	0.118
	(0.012)	(0.329)	(0.103)	
	[0.030]	[0.065]	[0.026]	
		[0.242]		
9 months	− 0.043	0.897	− 0.354	0.308
	(0.013)	(0.301)	(0.098)	
	[0.001]	[0.004]	[0.001]	
		[0.734]		
12 months	− 0.049	0.779	− 0.402	0.323
	(0.013)	(0.286)	(0.098)	
	[0.000]	[0.008]	[0.000]	
		[0.442]		

$\hat{\sigma}_t^2$ denotes the estimated conditional variance from the ARCH(2) model for $\Delta r_t^{(1)}$. The regressors are used as instruments. Newey–West standard errors in parentheses; p-values for tests of the hypothesis that the parameter is zero are in brackets. p-values for a test of the hypothesis that $\beta^{(N)}=1$ are in brackets and italics.

The results in Table 7 indicate that the use of the instrumental variables technique has an important impact on the estimates. In particular, and with the exception of the case of the 3-month rate, the estimated β parameters are numerically larger than those in Table 6 and generally more significant.[15] Moreover, the hypothesis that $\beta=1$ is not rejected. The δ parameters are all highly significant and also larger in absolute value than those in Table 6. Overall, these findings suggest that, in the presence of the term premium, term spreads are unbiased, but poor, predictors of future short-term rates.[16]

3.5. Parameter instability

The estimates discussed above are all obtained from data for the period 1992:01–2000:03 and implicitly assume that the parameters are stable in this period. However, following the onset of the Asian crisis in the summer of 1997, interest rates in Hong Kong experienced several episodes of sharp volatility. Moreover, the term structure has become

[15] It is possible that the insignificance of the parameters in the case of the 3-month rate is due to multicollinearity. The correlation between β and δ is -0.78 in this case, but between -0.72 and -0.61 in the other cases.

[16] To understand why they are poor predictors, recall that the parameters in Table 6 are obtained by exploiting the same orthogonality conditions as OLS. While these are selected to minimise the variance of the errors, the parameters in Table 7 (which stem from applying instrumental variables) are not. Thus, the latter regressions explain at best an even lower fraction of the variance of the dependent variable than the estimates in Table 6.

Table 7
GMM estimates of $\frac{1}{N}\sum_{i=0}^{N-1}\left(r_{t+i}^{(1)}-r_t^{(1)}\right)=\alpha^{(N)}+\beta^{(N)}\left(r_t^{(N)}-r_t^{(1)}\right)+\delta^{(N)}\log\left(\hat{\varepsilon}_t^2\right)+v_t^{(N)}$

Maturity of long rate (N)	1992:01–2000:03		
	$\alpha^{(N)}$	$\beta^{(N)}$	$\delta^{(N)}\times 100$
3 months	− 0.021	0.562	− 0.157
	(0.013)	(0.340)	(0.098)
	[0.121]	[0.102]	[0.112]
		[0.201]	
6 months	− 0.051	0.876	− 0.377
	(0.015)	(0.387)	(0.113)
	[0.001]	[0.026]	[0.001]
		[0.749]	
9 months	− 0.073	1.123	− 0.540
	(0.013)	(0.314)	(0.096)
	[0.000]	[0.001]	[0.000]
		[0.697]	
12 months	− 0.079	0.913	− 0.584
	(0.015)	(0.318)	(0.101)
	[0.000]	[0.005]	[0.000]
		[0.785]	

$\hat{\varepsilon}_t$ denotes the fitted errors from the ARCH(2) model for $\Delta r_t^{(1)}$. A constant, the term spread and $\hat{\sigma}_t^2$ are used as instruments for volatility. Newey–West standard errors in parentheses. p-values for tests of the hypothesis that the parameter is zero are in brackets. p-values for a test of the hypothesis that $\beta^{(N)}=1$ are in brackets and italics.

more upward sloping, despite the fact that interest rates have been relative stable since late 1998. This change in the behaviour of interest rates raises the issue whether the parameters in the equations estimated above changed as a consequence of the financial market turbulence caused by the Asian crisis. Since a change in the parameters would invalidate the analysis above, we next test for evidence of a structural break in 1997:06.

To do so, we first explore whether the ARCH(2) model for the change of the 1-month rate is stable across sample periods. Allowing for a shift in all parameters in 1997:06, we obtain $p=23.0\%$. Allowing for a break solely in μ_0 yields $p=28.9\%$, while allowing for a break solely in γ_0 yields $p=64.1\%$. These tests all suggest that the ARCH model is stable.

Next, we define a dummy variable D which equals zero for the period 1992:01–1997:05 and unity otherwise, and estimate

$$\frac{1}{N}\sum_{i=0}^{N-1}\left(r_{t+i}^{(1)}-r_t^{(1)}\right)=\alpha^{(N)}+\beta^{(N)}\left(r_t^{(N)}-r_t^{(1)}\right)+\delta^{(N)}\log\left(\hat{\varepsilon}_t^2\right)+\ldots+D \times\left\{\alpha^{*(N)}+\beta^{*(N)}\left(r_t^{(N)}-r_t^{(1)}\right)+\delta^{*(N)}\log\left(\hat{\varepsilon}_t^2\right)\right\}+v_t^{(N)}. \quad (12)$$

We can then test for a structural change in the parameters in 1997:06 by performing a Wald test of the hypothesis that $\alpha^{*(N)}=\beta^{*(N)}=\delta^{*(N)}=0$. In Table 8, we report estimates of Eq. (12) and p-values for the above test, and p-values for the joint test that $\alpha^{*(N)}=\beta^{*(N)}=\delta^{*(N)}=0,\ \beta^{(N)}=1$. Note that in no case can we reject the hypothesis that the parameters are stable before and after the onset of the Asian crisis. In fact, all the α^*,

Table 8
GMM estimates of

$$\frac{1}{N}\sum_{i=0}^{N-1}\left(r_{t+i}^{(1)}-r_t^{(1)}\right)=\alpha^{(N)}+\beta^{(N)}\left(r_t^{(N)}-r_t^{(1)}\right)+\delta^{(N)}\log\left(\hat{\varepsilon}_t^2\right)+\ldots$$
$$+D\times\left\{\alpha^{*(N)}+\beta^{*(N)}\left(r_t^{(N)}-r_t^{(1)}\right)+\delta^{*(N)}\log\left(\hat{\varepsilon}_t^2\right)\right\}+v_t^{(N)}$$

Maturity of long rate (N)	1992:01–2000:03					
	$\alpha^{(N)}$	$\beta^{(N)}$	$\delta^{(N)}\times 100$	$\alpha^{*(N)}$	$\beta^{*(N)}$	$\delta^{*(N)}\times 100$
3 months	− 0.026	1.175	− 0.183	− 0.001	− 0.737	− 0.043
	(0.013)	(0.446)	(0.098)	(0.023)	(0.611)	(0.178)
	[0.058]	[0.010]	[0.066]	[0.975]	[0.231]	[0.809]
	H_0: $\alpha^{*(N)}=\beta^{*(N)}=\delta^{*(N)}=0$, p-value = 0.1971					
	H_0: $\alpha^{*(N)}=\beta^{*(N)}=\delta^{*(N)}=0, \beta^{(N)}=1$, p-value = 0.160					
6 months	− 0.047	1.089	− 0.336	− 0.019	−0.270	− 0.203
	(0.028)	(0.414)	(0.211)	(0.035)	(0.624)	(0.270)
	[0.098]	[0.010]	[0.114]	[0.594]	[0.666]	[0.455]
	H_0: $\alpha^{*(N)}=\beta^{*(N)}=\delta^{*(N)}=0$, p-value = 0.538					
	H_0: $\alpha^{*(N)}=\beta^{*(N)}=\delta^{*(N)}=0, \beta^{(N)}=1$, p-value = 0.697					
9 months	− 0.092	1.659	− 0.658	0.014	− 0.701	0.047
	(0.049)	(0.646)	(0.359)	(0.051)	(0.746)	(0.371)
	[0.063]	[0.012]	[0.070]	[0.778]	[0.350]	[0.900]
	H_0: $\alpha^{*(N)}=\beta^{*(N)}=\delta^{*(N)}=0$, p-value = 0.643					
	H_0: $\alpha^{*(N)}=\beta^{*(N)}=\delta^{*(N)}=0, \beta^{(N)}=1$, p-value = 0.742					
12 months	− 0.133	1.603	− 0.947	0.053	− 0.862	0.320
	(0.076)	(0.727)	(0.542)	(0.078)	(0.802)	(0.522)
	[0.083]	[0.030]	[0.084]	[0.499]	[0.285]	[0.563]
	H_0: $\alpha^{*(N)}=\beta^{*(N)}=\delta^{*(N)}=0$, p-value = 0.626					
	H_0: $\alpha^{*(N)}=\beta^{*(N)}=\delta^{*(N)}=0, \beta^{(N)}=1$, p-value = 0.768					

$\hat{\varepsilon}_t$ denotes the fitted errors from the ARCH(2) model for $\Delta r_t^{(1)}$. D is a dummy which equals zero for the period 1992:01–1997:05 and unity otherwise. A constant, the term spread, $\hat{\sigma}_t^2$, the dummy, the dummy times the term spread, and the dummy times $\hat{\sigma}_t^2$ are used as instruments. Newey–West standard errors in parentheses. p-values for tests of the hypothesis that the parameter is zero are in brackets.

β^* and δ^* parameters are insignificantly different from zero, while the other parameters are similar to those reported in Table 7. The major exception is $\beta^{(3)}$, which is significant and close to unity. Overall, we interpret the results in Table 8 as not suggesting that a structural break took place following the onset of the Asian crisis.

4. Conclusions

In this paper, we have studied the term structure of interbank rates in Hong Kong, using data spanning 1992:01–2000:03. The main finding of the paper is that the term structure in Hong Kong does not seem to evolve over time in accordance with the simple EH. The reason for this appears to be the presence of time-varying term premia, which can be modelled quite well as depending on the volatility of innovations to the 1-month rate.

However, even if we incorporate term premia in the analysis, the forecasting ability of the term spreads remains very limited. Overall, the results thus suggest that term spreads are unbiased, but poor, predictors of future short-term rates in Hong Kong. However, to extract this information requires a model of the term premium. The difficulties in developing such a model coupled with the fact that it may change over time urges caution in using the slope of the term structure as a measure of interest rate expectations in Hong Kong.

Acknowledgements

I am grateful to the editor and an anonymous referee for constructive suggestions and to Priscilla Chiu, Petra Gerlach, Lok Sang Ho, Tony Latter and seminar participants at Lignan University and the HKIMR for useful comments. The views expressed in this paper are solely my own and not necessarily those of the HKMA, the HKIMR, its Council of Advisers or Board of Directors.

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