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**Estimates of time-varying term premia  
for New Zealand and Australia**

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## Estimates of time-varying term premia for New Zealand and Australia

### Abstract<sup>1</sup>

Forward rates in the money market are systematically higher than realised spot rates, reflecting an unobservable term premium. This paper uses a Kalman filter specification to produce time-varying estimates of the term premia in New Zealand and Australia. Three time series specifications are used to examine the properties of the premia, such as the average size, volatility, and the degree of mean reversion.

Compared to the constant term premia estimates, the time-varying estimates explain significantly more of the difference between forward and spot rates. The results suggest that the premium in New Zealand is slowly mean-reverting, while the Australian premium reverts quickly to the mean.

It is not clear whether the method of monetary policy implementation affects the term premium, although in New Zealand the premium has been smaller and less variable since the introduction of the Official Cash Rate in March 1999. A related finding is that the size of the term premium is correlated with the volatility of short-term rates.

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## 1 Introduction

Short-term interest rate expectations are of interest to market participants, including central banks, for a variety of reasons (Krippner and Gordon, 2001). Market interest rates provide the best measure of average expectations, but they first need to be disentangled from the term premium, which provides compensation for the risk borne in lending over longer periods of time. The term premium is not directly observable, so deriving a suitable estimate is necessary for examining interest rate expectations.

The process becomes slightly more straightforward when the central bank uses the cash rate as its main monetary policy tool. It is now common practice among the major central banks to review the cash rate at scheduled intervals, and to adjust it in increments of 25 basis points when necessary. The Reserve Bank of Australia (RBA) has taken this approach since January 1990, while the Reserve Bank of New Zealand (RBNZ) adopted this method in March 1999.<sup>2</sup> Experience has shown that a cash rate regime significantly reduces the amount of ‘noise’ in short-term market interest rates (Brookes and Hampton, 2000). As a result, changes in market interest rates are more likely to reflect changes in the market forecast of the policy rate.

Nevertheless, market interest rates are still highly volatile and the gap between forward and spot rates can be very large at times. Typically, variations in this gap are treated as one-for-one changes in market forecasts of the spot rate, and the term premium is assumed to be constant. But it is plausible that some of it is due to changes in the term premium as well – which raises the question of how to separate the two factors.

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<sup>2</sup> While it is common shorthand to refer to the central bank ‘setting’ the cash rate, in practice the method of policy implementation differs across countries. The RBA chooses a target cash rate, and aims for this target through its daily liquidity management operations (Rankin, 1992). The RBNZ’s Official Cash Rate (OCR) is not a target, but the mid-point of the rates at which it will borrow from or lend to banks in unlimited amounts. The RBNZ acts as the marginal supplier of liquidity, and leaves the banks to determine the market cash rate (Archer, Brookes and Reddell, 1999).

One econometric approach is to use an adaptive filtering technique, such as the Kalman filter, which allows parameters to vary stochastically over time. The advantage of the Kalman filter is that it does not require a prior specification of which factors affect the term premium, but instead relies on the properties of the observed data, along with some identifying assumptions. This feature is particularly useful, because the factors that may influence the term premium – such as interest rate volatility, credit risk, market liquidity, inflation uncertainty and political risk – cannot be measured directly, and are difficult to proxy with other variables.<sup>3</sup>

In this paper I use the methodology of Gravelle and Morley (2003) to produce time-varying estimates of the term premia for New Zealand and Australian short-term rates. As well as producing constant term premia estimates as a base case, I consider two alternative specifications – a mean-reverting specification, and a random walk that allows for the possibility of permanent shocks to the term premia.

I focus on Australia as well as New Zealand, to contrast the changes in the method of implementing monetary policy in New Zealand. The RBA used the overnight cash rate as its main policy tool during the entire period I have examined here, while the RBNZ adopted a cash rate regime about halfway through the sample period. The structures of the Australasian money markets are similar in most other respects, so the broad trends of the term premia is likely to reflect country-specific factors.

The paper proceeds as follows: Section 2 reviews some of the previous work on both constant and time-varying estimates of the term premium. Section 3 provides brief details of the model and the properties of the data. Section 4 presents the key results. Section 5 discusses some of the features of the results, and explores some uses for the output of the model. Section 6 summarises the findings, and notes some avenues for further research.

<sup>3</sup> Note that the term premium is different from the concept of the ‘country risk.’ premium, which would be reflected in both the spot and forward rates.

## 2 Review

A forward interest rate can always be decomposed as follows:

$$f_{t,j} = E_t[r_{t+j}] + \alpha_{t,j} \quad (1)$$

$f_{t,j}$  is the forward interest rate with settlement  $j$  periods ahead, and  $E_t[r_{t+j}]$  is the average market expectation of the  $j$ -period-ahead spot rate, given the information available at time  $t$ .  $\alpha_{t,j}$  is the systematic difference between the spot and forward rates, known as the ‘term premium’ because it is associated with the term to maturity.

Traditional tests for market efficiency make three assumptions about this relationship. First, interest rate expectations are fulfilled on average, so that  $E_t[r_{t+j}] = r_{t+j}$ . Second, the term premium  $\alpha_{t,j}$  is fixed over time. Third, there are no systematic forecast errors, so that  $e_{t+j}$  is white noise with zero mean. Most studies test these assumptions, and estimate the size of the term premium, using a linear regression of the following relationship:

$$f_{t,j} = \alpha_j + \lambda_j r_{t+j} + e_{t+j} \quad (2)$$

The parameter  $\lambda_j$  will be equal to one in an efficient market, and it is sometimes imposed as one to simplify the estimation. When  $\lambda_j$  is found to be different from one, it is usually treated as a rejection of market efficiency. Notably, there have been numerous studies, particularly for the US, that have found  $\lambda_j$  to be significantly less than one<sup>4</sup>. Taken at face value, this might suggest that forward rates are biased predictors of spot rates.

Mankiw and Miron (1986) note that a time-varying term premium could explain these results. A linear regression of equation (2) imposes a constant value for the term premium, and if the true premium actually varies over time, the estimates of  $\lambda_j$  can be biased downward. This means that the findings of a bias in forward rates

<sup>4</sup> See Cook and Hahn (1990), for example, for a review of the post-war US literature.

may be due to the estimation method, rather than a feature of the market.

With this in mind, some studies have attempted to produce time-varying estimates of the term premium. Finance theory offers no clues as to the correct specification for a time-varying premium, so these studies have focused on factors that could plausibly influence the premium, such as the volatility of key financial and economic variables. For example, Engle, Lilien and Robins (1987) developed the ARCH-M approach, where the size of the term premium is influenced by the conditional variance of interest rates. Lee (1995) uses ARCH measures of the conditional variances of macroeconomic variables.

Some recent studies have used the Kalman filter to produce time-varying term premia estimates based on the properties of the data. Iyer (1997) uses this approach for US short-term rates and finds significant variance in the term premium. Bhar (1993) also finds a time-varying premium in Australian rates, albeit from a short and atypical sample period. Gravelle and Morley (2003) examine Canadian forward rates and find support for a term premium that is time-varying but stationary over the long term.

Studies of the New Zealand and Australian markets have generally found that  $\lambda_j$  is not significantly different from one, and hence they have focused on the size of  $\alpha_j$ . Guthrie, Wright and Yu (1999) and Krippner (2002) found a statistically significant term premium on money market rates. Petro, McDermott and Tripe (2003) found a premium of more than 100 basis points on bank bill rates, which fell significantly after the introduction of the OCR. For Australia, Lowe (1995) found a premium of about 40 basis points on six-month bills, while Gordon (2002) found a premium of about 10 basis points for the same horizon.

None of these studies have addressed the question of a time-varying premium in any detail, although the results suggest that a constant term premium is a reasonable assumption. Even so, the point estimates of the premium tend to be very imprecise, and there appears to be some scope for improving the quality of the estimates by allowing for the possibility of a time-varying premium.

### 3 Model

#### 3.1 Specification

Subtracting the realised spot rate  $r_{t+j}$  from both sides of equation (1) gives:

$$efr_{t,j} = \alpha_{t,j} + u_{t+j} \quad (3)$$

Where  $efr_{t,j} = f_{t,j} - r_{t+j}$  is the  $j$ -period ahead excess forward return, and the market forecast error  $u_t = E[r_{t+j}] - r_{t+j}$ . Expectations are realised on average over the long term, so the forecast error in equation (3) has an expected value of zero. The forecast error is also assumed to be uncorrelated with market interest rates at time  $t$ . This corresponds to the idea that financial market expectations are rational (Muth, 1961).

I tested three different specifications for the unobserved term premium. The first is the base case of a ‘‘constant’’ specification, which reduces the model to a linear estimation of equation (3):

$$\alpha_{t,j} = \bar{\alpha}_j \quad (4)$$

The second is the ‘‘non-stationary’’ specification, where the term premium is assumed to follow a random walk, with no time trend:

$$\alpha_{t,j} = \alpha_{t-1,j} + v_t \quad (5)$$

The final specification is ‘‘mean-reverting’’, where the term premium follows a stationary first-order autoregressive process:

$$\alpha_{t,j} = c + \phi\alpha_{t-1,j} + v_t \quad (6)$$

Where  $\phi$  is less than one in absolute value. Note that the first two specifications are restrictions of the third; the constant scenario restricts  $\phi$  and the variance of  $v_t$  to zero, while in the non-stationary case the restrictions are  $c = 0$  and  $\phi = 1$ .

The question of whether the term premium is stationary is more of an empirical than a theoretical one. Some evidence suggests that real interest rates are mean-reverting over very long periods, but the sample period used here is probably too small to produce strong conclusions. The relevant issue is the degree of mean reversion – that is, the relative weighting of transitory and permanent shocks to the term premium. The non-stationary specification attributes a larger portion of high-frequency changes in excess returns to changes in the term premium, so it should provide the strongest evidence against the assumption of a constant premium. On the other hand, the mean-reverting specification is the least restrictive, and hence it should better reflect the ‘true’ premium.

The final issue is the specification of the errors. The error term  $v_t$  represents the shocks to the process that generates the term premium, which can arise from any of the factors mentioned earlier such as changes in credit risk or market liquidity. These shocks are assumed to occur at random, with no serial correlation.

The market forecast error  $u_{t+j}$  is assumed to follow a moving average process with  $j-1$  lags, as in Morley (1999). This reflects the fact that new information arrives between the date the forecasts (and the term premia) are determined, and the date the cash rate is announced. Individual pieces of new information are independently distributed, and the variance of the shocks is assumed to be constant over time, but they have a cumulative impact on the forecast error for horizons more than one period ahead:

$$u_{t+j} = e_{t+j} + \theta_1 e_{t+j-1} + \theta_2 e_{t+j-2} + \dots + \theta_{j-1} e_{t+1} \quad (7)$$

$$e_{t+j} \sim N(0, \sigma_e^2) \quad (8)$$

The model can now be written in state-space form.<sup>5</sup> To produce estimates of the term premium, the model requires up to four parameters, depending on the specification – the mean  $c$ , the rate of mean reversion  $\phi$ , and the variances of the error terms  $\sigma_v$  and  $\sigma_e$  (plus

the coefficients for the moving-average error terms). I used the OPTMUM optimisation routine in the GAUSS econometric program to derive the maximum likelihood estimates of the parameters, and the corresponding term premia estimates, for each specification. Numerical derivatives were used to calculate the asymptotic standard errors of the parameters. The error variance parameters  $\sigma_v$  and  $\sigma_e$  were constrained to be positive, and for the mean-reverting case the parameter  $\phi$  was constrained to be less than one in absolute value. In most cases the results were robust to a range of initial parameter settings; I discuss the exceptions later in the paper.

### 3.2 Data

Market interest rates are taken from the bank risk curve – namely, bank bills and forward rate agreements for the forward interest rates, and the inter-bank cash rate for the spot rate. Bank rates are more suitable than government risk (Treasury bill) rates, because the market is more active and there is a more complete range of maturities. The Appendix provides details on the data sources and the construction of the data.

Table 1 summarises the features of the excess forward returns (the difference between the spot and forward rates) for both countries. The first point to note is that the mean increases with the time horizon, which is consistent with a term premium that also increases with the horizon. The variation in excess returns also increases with the forward horizon, although this may reflect larger forecast errors at longer horizons, rather than greater variance in the term premium.

<sup>5</sup> Full details of the state-space formulation are available in Gravelle and Morley (2003). For a more general overview of state-space models and the Kalman filter, see Basdevant (2003).

**Table 1**  
**Descriptive statistics for excess forward returns**

	New Zealand			Australia		
	<i>Mean</i>	<i>Std Dev</i>	<i>ADF*</i>	<i>Mean</i>	<i>Std Dev</i>	<i>ADF*</i>
<i>j=1</i>	0.201	0.337	-2.94	0.082	0.117	-4.03
<i>j=2</i>	0.243	0.575	-3.04	0.084	0.200	-3.88
<i>j=3</i>	0.292	0.787	-2.62	0.118	0.293	-3.61
<i>j=4</i>	0.270	0.973	-3.58	0.123	0.398	-5.09
<i>j=5</i>	0.299	1.133	-4.55	0.189	0.499	-3.47
<i>j=6</i>	0.330	1.258	-3.74	0.195	0.573	-3.85
<i>j=7</i>	0.416	1.376	-3.65	0.352	0.717	-3.18
<i>j=8</i>	0.386	1.385	-3.40	0.407	0.724	-3.49
<i>j=9</i>	0.471	1.516	-3.48	0.455	0.771	-4.16
<i>j=10</i>	0.569	1.533	-3.84	0.462	0.921	-3.79
<i>j=11</i>	0.587	1.550	-3.78	0.623	0.915	-3.05
<i>j=12</i>	0.655	1.561	-3.25	0.676	0.949	-3.65

\* The optimal numbers of lags were selected using the Schwartz information criteria. Critical values are -2.89 at the 5% level and -2.58 at the 10% level.

Table 1 also gives the results of the augmented Dickey-Fuller test for a unit root in excess forward returns. While the presence of a unit root is not a problem for Kalman filter estimation, the test provides useful information about the nature of the term premium. Specifically, if the excess returns are stationary, then the term premium is also likely to be stationary. At every horizon, the test rejects the hypothesis of a unit root at the 5 per cent level of significance (except at the three-month horizon for New Zealand, where it rejects a unit root at the 10 per cent level). However, as noted before, it is difficult to draw strong conclusions from this small sample.

As an aside, Johansen cointegration tests confirm the findings of Krippner (2002) and Gordon (2002) that there is a long-run relationship between the cash rate and forward rates. The tests found a single cointegrating vector at each horizon, and the standardised coefficients for the cash rate, equivalent to  $\lambda_j$  in equation (2), were all close to one.

**Figure 1**  
**Excess forward returns for New Zealand and Australia**

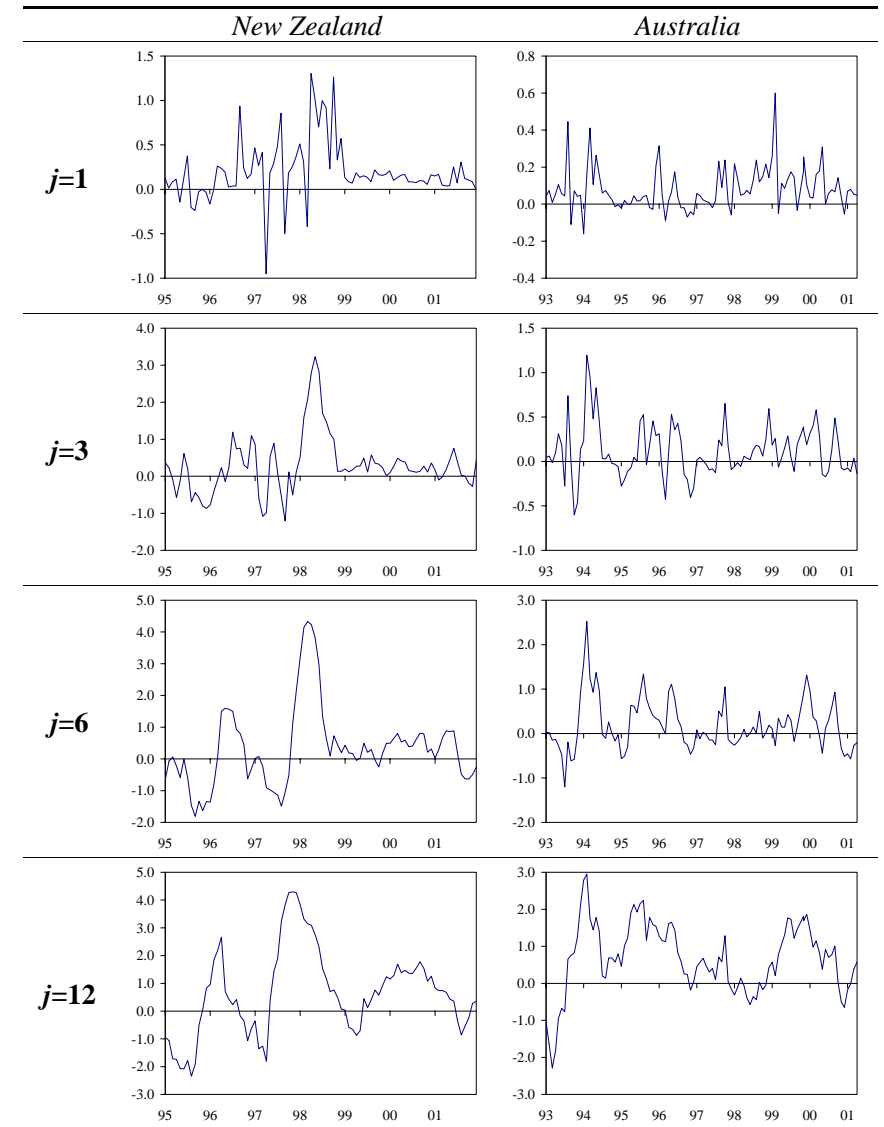


Figure 1 shows the excess forward returns for New Zealand and Australia.<sup>6</sup> The forward returns become more persistent as the horizon increases, and are noticeably cyclical at the twelve-month horizon. This is consistent with the idea that shocks to interest rate expectations have a cumulative effect at longer horizons, creating serial correlation in the excess forward returns. The excess returns for New Zealand are noticeably less volatile after March 1999, when the RBNZ adopted a cash rate regime. In comparison, the volatility of the excess returns in Australia is fairly consistent over the sample period.

## 4 Results

### 4.1 Constant specification

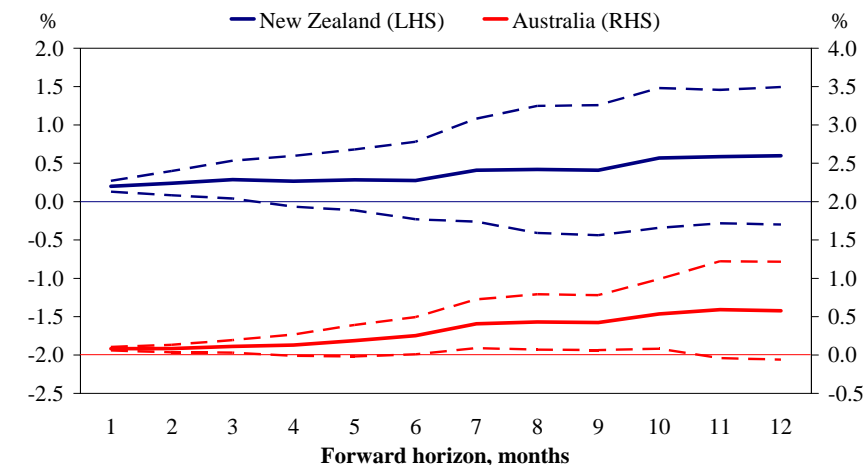
Figure 2 presents the profiles of the constant term premia estimates for New Zealand and Australia, and the 95 per cent confidence intervals. The estimates fit with the traditional view of the term premium, in that they are positive at every forward horizon, and tend to increase with the horizon. The cross-country estimates are very similar for horizons longer than six months, but the Australian estimates for up to three-month horizons are substantially lower. The estimates for both countries have very wide confidence intervals at every horizon. In fact, the estimates for New Zealand are not significantly different from zero at most horizons. The Australian estimates have slightly smaller confidence intervals, and are significantly greater than zero at most horizons.

The premium at the one-year horizon is about 60 basis points in both countries, which is slightly larger than the estimates in Krippner (2002) and Gordon (2002). These studies examined the period from 1999 onward, which did not cover a full monetary policy cycle. Anecdotes from dealers in the New Zealand market indicate that their perceptions of the term premium change over the monetary policy cycle, although not necessarily in tandem with the cycle.<sup>7</sup>

<sup>6</sup> Figure 1 presents only a selection of time horizons, for the sake of space, but the broad trends in the data for the other horizons are similar to those shown here.

<sup>7</sup> For example, the term premium tends to be larger when there is greater uncertainty about the direction of monetary policy. When dealers expect the cash

**Figure 2**  
**Constant term premia estimates**



Dashed lines are 95% confidence intervals.

### 4.2 Non-stationary and mean-reverting specifications

Table 2 compares the log-likelihood values for each specification. The time-varying specification naturally has greater explanatory power than the constant specification, because it has fewer restrictions. I use the Wald test to determine whether the time-varying specification is a significant improvement on the constant specification.

For New Zealand, the non-stationary specification is an improvement for seven of the twelve horizons, while the mean-reverting specification is significantly better for eleven horizons. Garbade (1977) notes that this test tends to understate the significance of time-varying parameters, because the distribution under the null hypothesis (constant parameters) is more concentrated

rate to remain unchanged for a long time, they tend to buy and hold bank bills, so that the term premium is bid down. Of course this is anecdotal, and could be offset by other factors.

than the  $\chi^2$  distribution used in the Wald test. This means that the time-varying estimates are probably even more significant than the tests indicate.

The Wald test finds little to differentiate between the mean-reverting and non-stationary specifications. In any case, Dickey and Fuller (1981) warn that this test is biased when the null hypothesis is for non-stationarity. Since both sets of time-varying estimates are superior to the constant estimates, it appears that the results do not hinge on the correct specification of the degree of mean reversion.

**Table 2**  
**Log-likelihood values for New Zealand and Australia**

Forward horizon	New Zealand			Australia		
	Constant	Non-stationary	Mean-reverting	Constant	Non-stationary	Mean-reverting
$j=1$	-32.41	<b>-27.86</b>	<b>-25.59</b>	-72.54	<b>-68.45</b>	-72.52
$j=2$	-63.77	<b>-61.62</b>	<b>-58.00</b>	-26.28	<b>-21.32</b>	-26.28
$j=3$	-60.37	<b>-57.66</b>	<b>-54.47</b>	-6.64	-6.12	-6.61
$j=4$	-65.30	<b>-63.37</b>	<b>-60.66</b>	-23.29	-23.23	-23.29
$j=5$	-63.37	<b>-61.19</b>	<b>-59.13</b>	-31.66	-31.21	-31.66
$j=6$	-56.70	-55.01	<b>-52.08</b>	-44.93	-44.32	-44.93
$j=7$	-69.02	-67.39	<b>-66.01</b>	-55.44	-54.52	-55.10
$j=8$	-69.98	-68.13	<b>-66.58</b>	-50.85	-49.82	-50.85
$j=9$	-59.95	-58.08	<b>-55.92</b>	-53.21	-52.20	-53.21
$j=10$	-66.83	<b>-64.88</b>	<b>-61.83</b>	-68.05	-66.80	-68.05
$j=11$	-56.48	-54.58	-53.69	-53.82	-53.33	-53.82
$j=12$	-61.80	<b>-59.81</b>	<b>-57.48</b>	-48.23	-46.65	-48.23

Wald test scores are calculated as  $2*(LLR_{unrestricted} - LLR_{restricted})$ . The test statistics for the non-stationary specification are distributed as  $\chi^2(1)$  with a 5% critical value of 3.84. The mean-reverting statistics are distributed as  $\chi^2(2)$  with a 5% critical value of 5.99. Figures in bold indicates that the estimates are more significant than the constant premium estimates at the 5% level.

In contrast, the Australian results provide only slight support for a time-varying premium. The non-stationary specification produces higher log-likelihood values, but they are only significant for the first two horizons. The mean-reverting specification performed even worse – the parameter  $\phi$  converged to zero at every horizon, so that the estimated term premia became a constant value with random disturbances in each period. Since these disturbances have no predictive power, the log-likelihood values for the mean-reverting results are identical to the constant premium estimates.

Recognising that maximum likelihood estimation is prone to having several local maxima, I re-estimated the model for Australia using a range of initial parameter settings. The results were robust for horizons up to eight months, but for longer horizons it was possible to produce estimates for  $\phi$  of around 0.8, similar to the results for New Zealand. This finding is consistent with Figure 2, which shows a lack of persistence in the Australian excess returns, except at the longer horizons. For the sake of consistency, the estimates presented from here on are those with  $\phi = 0$  at every horizon.

Table 3 presents the optimised parameters for the New Zealand data under the mean-reverting specification. The variances of the term premia shocks  $\sigma_v$  are significantly greater than zero at every horizon, which is consistent with a time-varying term premium. The mean-reversion parameters range from 0.7 to 0.9 at each horizon, and while they are significantly greater than zero, they are not significantly different from one. As noted before, there is little to distinguish between the mean-reverting and non-stationary specifications.

The results in table 3 show that the mean-reverting estimates are economically as well as statistically significant. On average, the standard error of shocks to the term premium  $\sigma_v$  is about the same magnitude as for the market forecast error  $\sigma_e$ . In other words, changes in the term premium explain about half of the variation in excess returns. This ratio is comparable to those found in earlier studies. Gravelle and Morley (2003) calculated ratios of 30 to 80 per cent for Canada using the same methodology, while MacDonald and MacMillan (1994) found a ratio of 30 to 60 per cent for the UK based on surveys of economists' forecasts.



Table 4 shows the parameters for the mean-reverting specification for Australia. As noted earlier, the mean reversion parameter  $\phi$  is virtually zero at each horizon. The variance of the term premia shocks are small compared to New Zealand, but are still significantly greater than zero at every horizon. On average the term premium contributes about 30 per cent of the volatility in excess forward returns.

Figure 3 shows the profiles of the non-stationary and mean-reverting term premia estimates for New Zealand. (The first 12 months of results are not shown because the confidence intervals tended to be extremely wide for this period.) An unusual feature of these estimates is that they reach more extreme values at the shorter horizons. One possible explanation for this comes from Gurkaynak, Sack and Swanson (2002), which shows that a significant portion of the market impact of a monetary policy surprise is due to the timing rather than the size of the shock. Over longer horizons, the market is fairly accurate at predicting the total amount of easing or tightening, but the timing of the changes is a significant source of forecast error. However, this is not a feature of the Australian estimates, so it could also reflect the operation of monetary policy in New Zealand – a theme explored in the next section of the paper.

Figure 4 presents the Australian term premia estimates for both time-varying specifications. The non-stationary estimates are much less variable than their New Zealand counterparts, and the confidence intervals are wide enough to allow for the possibility that the term premium is actually constant over time. Although the non-stationary specification did not include a time trend, the estimates do seem to have a slight downward trend over the sample period. The mean-reverting estimates, as noted earlier, are essentially flat with random disturbances in each period. These disturbances tend to be smaller at longer horizons.

**Table 3**  
**New Zealand parameter estimates – mean-reverting**

<i>Forward horizon</i>	$\sigma_e$	$\sigma_v$	$\phi$	$c$
$j=1$	<b>0.282</b> (0.027)	<b>0.093</b> (0.041)	<b>0.857</b> (0.106)	<b>0.210</b> (0.080)
$j=2$	<b>0.323</b> (0.143)	<b>0.323</b> (0.165)	<b>0.708</b> (0.204)	0.235 (0.132)
$j=3$	<b>0.249</b> (0.099)	<b>0.223</b> (0.066)	<b>0.836</b> (0.074)	0.279 (0.236)
$j=4$	<b>0.217</b> (0.087)	<b>0.343</b> (0.116)	<b>0.854</b> (0.075)	0.262 (0.294)
$j=5$	<b>0.397</b> (0.070)	<b>0.186</b> (0.057)	<b>0.889</b> (0.083)	0.374 (0.291)
$j=6$	0.225 (0.321)	<b>0.278</b> (0.096)	<b>0.833</b> (0.134)	0.427 (0.317)
$j=7$	<b>0.389</b> (0.163)	<b>0.319</b> (0.112)	<b>0.848</b> (0.127)	0.445 (0.361)
$j=8$	<b>0.410</b> (0.179)	<b>0.302</b> (0.120)	<b>0.873</b> (0.094)	0.525 (0.408)
$j=9$	0.368 (0.385)	<b>0.231</b> (0.066)	<b>0.890</b> (0.079)	0.647 (0.430)
$j=10$	<b>0.366</b> (0.161)	<b>0.297</b> (0.099)	<b>0.876</b> (0.101)	0.662 (0.449)
$j=11$	<b>0.394</b> (0.105)	<b>0.265</b> (0.094)	<b>0.907</b> (0.055)	0.818 (0.446)
$j=12$	0.161 (0.158)	<b>0.404</b> (0.122)	<b>0.907</b> (0.050)	0.893 (0.539)

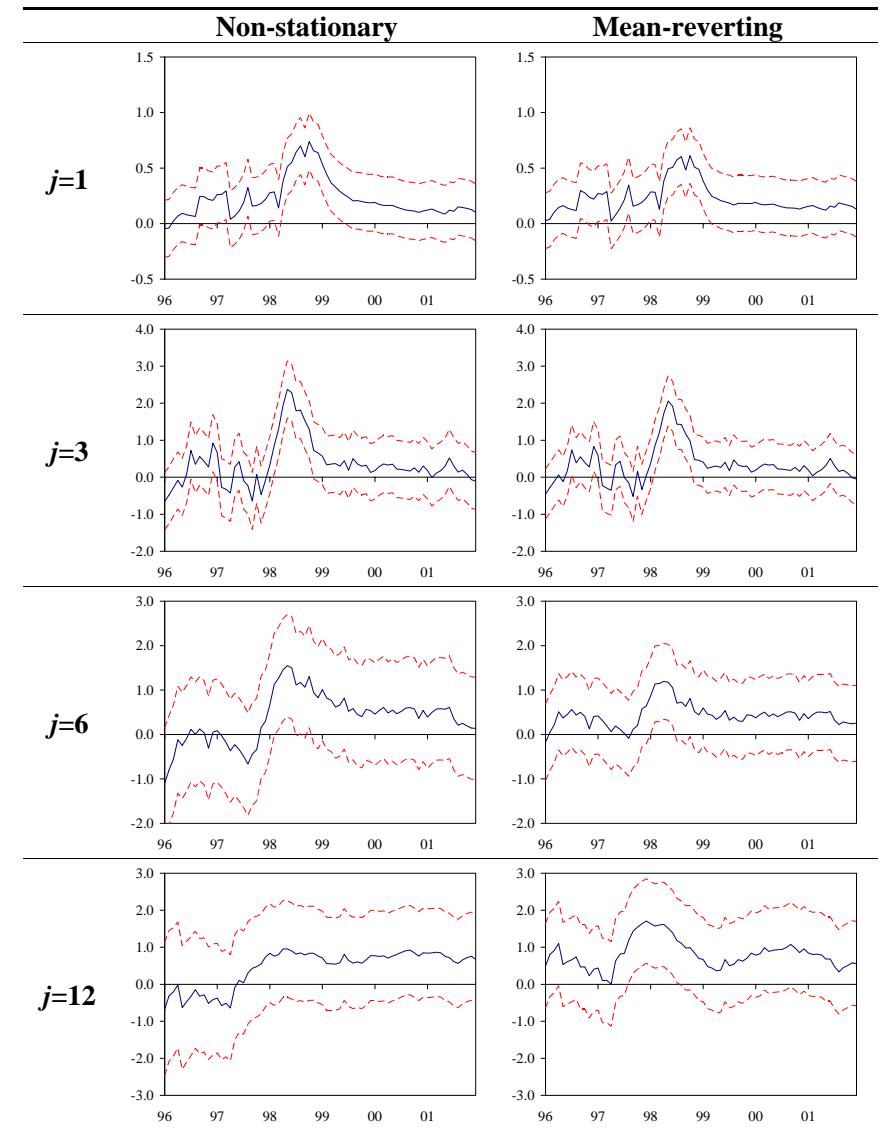
Standard errors are in parentheses. Figures in bold are significantly greater than zero at the 5% level.

**Table 4**  
**Australian parameter estimates – mean-reverting**

<i>Forward horizon</i>	$\sigma_e$	$\sigma_v$	$\phi$	$c$
$j=1$	<b>0.078</b> (0.034)	<b>0.043</b> (0.008)	0.003 (0.013)	<b>0.082</b> (0.012)
$j=2$	<b>0.074</b> (0.008)	<b>0.039</b> (0.002)	0.000 (0.002)	<b>0.084</b> (0.025)
$j=3$	<b>0.108</b> (0.033)	<b>0.025</b> (0.002)	0.000 (0.025)	<b>0.112</b> (0.042)
$j=4$	<b>0.053</b> (0.024)	<b>0.034</b> (0.007)	0.000 (0.001)	<b>0.142</b> (0.065)
$j=5$	<b>0.131</b> (0.047)	<b>0.035</b> (0.004)	0.000 (0.002)	<b>0.204</b> (0.089)
$j=6$	0.112 (0.098)	<b>0.044</b> (0.008)	0.000 (0.003)	<b>0.268</b> (0.091)
$j=7$	<b>0.151</b> (0.077)	<b>0.069</b> (0.004)	0.000 (0.01)	<b>0.405</b> (0.069)
$j=8$	<b>0.121</b> (0.048)	<b>0.053</b> (0.003)	0.000 (0.002)	<b>0.423</b> (0.054)
$j=9$	<b>0.103</b> (0.036)	<b>0.063</b> (0.002)	0.000 (0.007)	<b>0.436</b> (0.067)
$j=10$	<b>0.189</b> (0.068)	<b>0.036</b> (0.001)	0.000 (0.009)	<b>0.528</b> (0.084)
$j=11$	0.096 (0.059)	<b>0.035</b> (0.017)	0.000 (0.010)	<b>0.530</b> (0.083)
$j=12$	<b>0.137</b> (0.044)	<b>0.022</b> (0.008)	0.000 (0.011)	<b>0.527</b> (0.091)

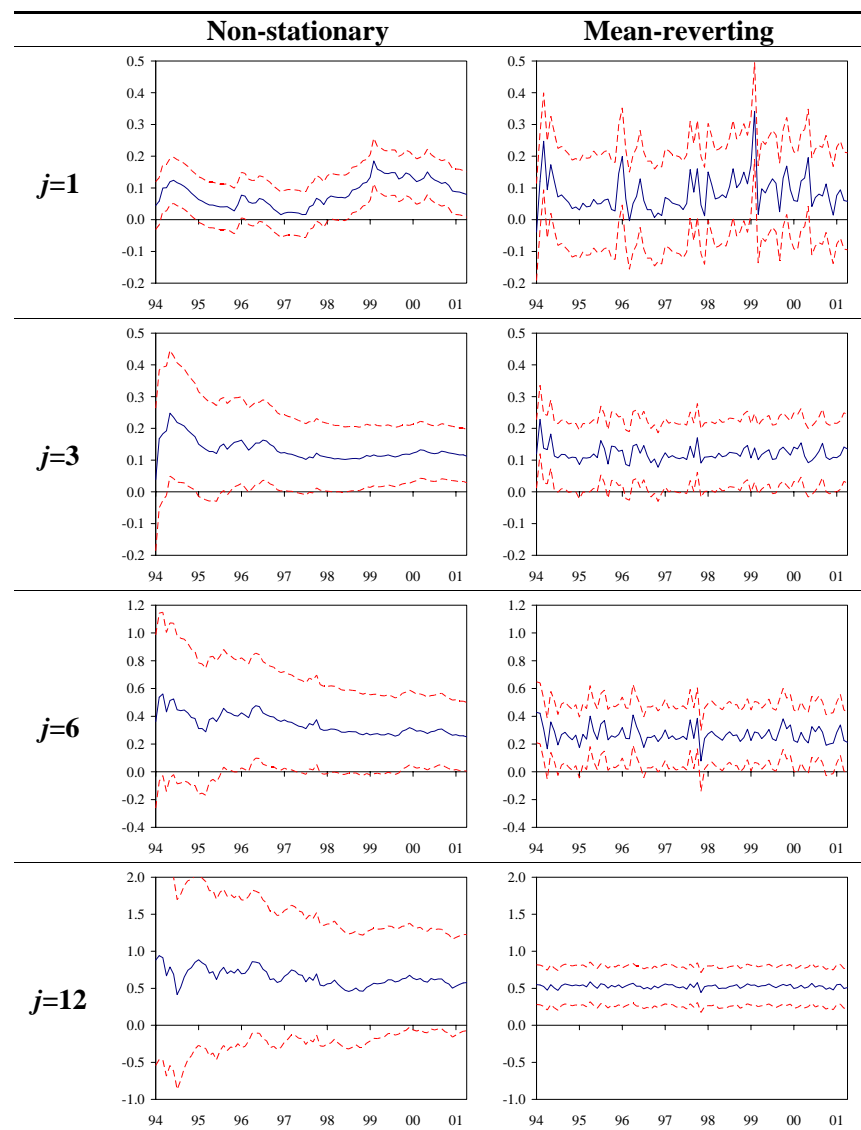
Standard errors are in parentheses. Figures in bold are significantly greater than zero at the 5% level.

**Figure 3**  
**Term premia estimates for New Zealand**



The dashed lines are 95% confidence intervals.

**Figure 4**  
**Term premia estimates for Australia**



The dashed lines are 95% confidence intervals.

## 5 Analysis

### 5.1 Method of monetary policy implementation

The process of financial market deregulation in the 1980s culminated with the Reserve Bank Act in 1989. The Act limits the role of monetary policy to achieving price stability – the definition of which has been refined over the years, through a series of Policy Targets Agreements signed by the RBNZ Governor and the Minister of Finance. However, the RBNZ has operational independence and is free to choose the methods it uses to achieve this goal. As a result, the implementation of monetary policy has evolved over time, and the second half of the 1990's was a period of dramatic change.

The RBNZ initially targeted quantity-based measures of monetary conditions, such as the level of settlement cash. However, its most effective tool was 'moral suasion' through published forecasts, including key assumptions about the future paths of the exchange rate and short-term interest rates, to guide the market towards the desired monetary settings. In most instances the market automatically adjusted quantities and prices to more desirable levels, and intervention was only occasionally required (Huxford and Reddell, 1996).

The RBNZ has long recognised that the exchange rate has a strong influence on output and prices, due to New Zealand's open economy and reliance on overseas trade, and hence it is a major determinant of overall monetary conditions.<sup>8</sup> In mid-1997 the RBNZ formally acknowledged the importance of this relationship by introducing the Monetary Conditions Index (MCI). The index was calculated so that a 2 per cent fall in the trade-weighted exchange rate was equivalent to a 1 percentage point fall in the three-month bank bill rate, in terms of the expected impact on inflation. It was intended to be a communication tool, to quantify the vague concept of the 'stance' of monetary conditions.

<sup>8</sup> It is important to note that although the RBNZ monitors the exchange rate as well as interest rates, it only implements monetary policy through the latter market. The RBNZ has not intervened in the foreign exchange market since the currency was floated in 1985.

During this time, public statements on policy settings were made irregularly but relatively frequently, and for the most part they were in response to fluctuations in the inherently volatile exchange rate (Brookes and Hampton, 2000). The emphasis on allowing the market to self-adjust also meant that, at times, market interest rates tracked the day-to-day and even intra-day volatility of the exchange rate.

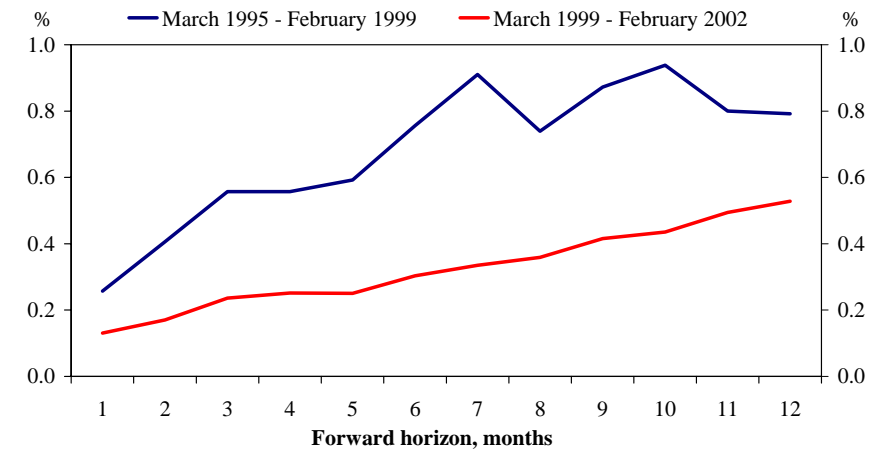
In March 1999, the RBNZ switched to a price-based approach to policy setting, and the Official Cash Rate (OCR) became the sole instrument for setting and communicating the stance of monetary policy. In addition, decisions about changes in policy settings were limited to eight scheduled dates per year, in normal circumstances.<sup>9</sup>

The change to a cash rate regime could have affected the term premium in two ways. First, the new regime greatly reduced the amount of day-to-day volatility in interest rates. From July 1997 to February 1999, the average daily change in the market cash rate was 16 basis points; from March 1999 it fell to almost zero (Brookes and Hampton, 2000). In the pre-OCR period, the degree of ambiguity about the ‘true’ level of the cash rate may have led investors to demand a larger premium for holding bank bills. If so, the mean of the term premia estimates should have fallen since March 1999. The second effect is that there should be much less variation in both the term premium and in market forecast errors during the OCR period.

Figure 5 shows the constant term premia estimates after splitting the sample into the pre-OCR and OCR regime periods. At every horizon, the term premium is about 25-50 basis points lower during the OCR period compared to the pre-OCR period. This difference is very important in policy-setting terms – it is equal to one or two increments in the OCR. The variance of the forecast errors is also much lower in the OCR period.

<sup>9</sup> The only unscheduled change so far was on 19 September 2001, following the terrorist attacks in the US.

**Figure 5**  
**New Zealand’s term premium before and after OCR**



However, there are two reasons to treat this result with caution. First, the difference between the two periods is not statistically significant due to the wide confidence intervals. Second, the Australian term premium also fell by about 20 basis points in the same period, so it is possible that this was part of a wider trend towards lower term premia in recent years.

The mean-reverting estimates for the OCR period converged to a constant value. This could be a misleading result due to the small sample size (36 observations), but it is consistent with the smaller variance in excess forward returns since March 1999. It is also consistent with the estimates for Australia, where the cash rate was used over the entire sample period.

Overall, this does not provide satisfactory evidence that the method of monetary policy implementation affected the behaviour of the term premium. It is easy to overstate the impact of the MCI, because it was introduced during an unusually volatile period for financial markets worldwide. The Asian currency crisis began in 1997 and continued throughout the following year, and the second half of 1998 saw the Russian debt default and the collapse of Long Term Capital Management. On top of this, the New Zealand economy had

to contend with a severe drought, and the exchange rate was falling rapidly from levels that most observers considered to be overvalued.

## 5.2 Negative term premia

Figure 3 suggests that the New Zealand term premia were often negative between 1996 and 1997, more so for the non-stationary specification. Gravelle and Morley (2003) also found negative term premia for Canada in the late 1980's, and suggested this was due to a "peso problem", as discussed in Bekaert, Hodrick and Marshall (1997). Early on in an inflation-fighting regime, the market may be sceptical about the central bank's commitment to the policy, and may continue to anticipate a return to more expansive monetary conditions. As a result, markets can under-predict short-term rates for extended periods, but can still be considered rational given the information available at the time.

While inflation targeting was well established in New Zealand by 1995, the market had some concerns about the way the RBNZ pursued this policy. At the time, New Zealand's interest rates were high relative to the rest of the world, and the market believed that this margin was unsustainable. The RBNZ felt that higher rates were appropriate, due to higher than expected growth, strong inward migration, and a surge in borrowing for housing. As a result, the money market yield curve was often downward-sloping (as was the longer-term government bond yield curve), and excess forward returns were persistently negative.

By similar reasoning, the large term premia estimates during 1998 could be described as a 'reverse peso problem': as the exchange rate fell, the market may have over-estimated the RBNZ's commitment to maintaining a certain level of the MCI, based on its earlier behaviour. Instead, the RBNZ allowed interest rates to fall sharply in the second half of 1998 – the three-month bank bill rate fell by over 300 basis points in the space of four months. In contrast, the RBA only reduced the cash rate by 25 basis points in the second half of 1998, which was less than the market expected. As a result, the excess forward returns for Australia were low or negative during this period.

## 5.3 Cross-sectional term premium function

In order to examine the behaviour of the term premium over time, it is useful to express it as a function of the time horizon. Previous studies have found that the term premium function increases monotonically with the forward horizon, at a decreasing rate (McCulloch, 1987). A single-parameter square-root function produces such a profile:

$$\alpha_{t,k} = \psi_t \sqrt{k} \quad (9)$$

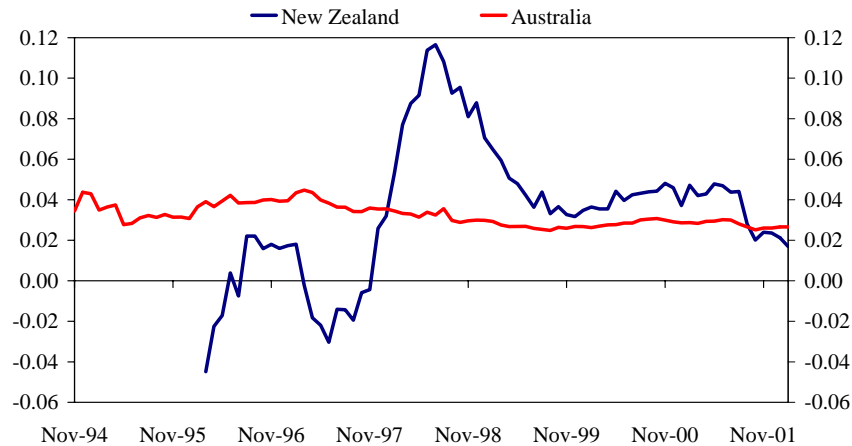
Where  $k$  is the forward time horizon in days, and is treated as the mid-point of the forward horizon, i.e. it is 15 days for  $j=1$ , 45 days for  $j=2$ , and so on. The function was fitted to the cross-section of the term premia estimates for each month. Figure 6 plots the time series of the slope parameter  $\psi_t$  for the non-stationary estimates, for New Zealand and Australia.<sup>10</sup>

The chart highlights the lack of a relationship between the term premia for New Zealand and Australia (in fact, the correlation is  $-0.45$ ). The New Zealand term premium has become more stable since 1999, while the Australian premium is just as variable. This supports the idea that changes in the term premia are largely due to country-specific factors.

Single-factor specifications of the term premium are common in the literature.<sup>11</sup> The major shortfall of this specification is that it assumes the shape of the term premium function is constant over time, and that only the slope of the curve changes. Figure 3 shows that this does not adequately describe the New Zealand data in the pre-OCR period, where the term premium appears to have been much larger for the three- to six-month horizons. However, this unusual result appears to be confined to the MCI period, and it is not obvious that the term premium function needs to incorporate this.

<sup>10</sup> A change of 0.01 in  $\psi_t$  is equal to a change in the term premium at the one-year horizon of about 19 basis points (the square root of 345 days).

<sup>11</sup> See, for example, Tzavalis and Wickens (1997).

**Figure 6****Term premium slope parameter  $\psi_t$** 

Ultimately, it may be more worthwhile to focus only on horizons up to six months ahead. Over the period examined here, the typical monetary tightening or easing period has been less than a year. One-year-ahead forecasts are unlikely to be very meaningful, since there is a non-trivial chance that monetary policy will change direction within that time.

#### 5.4 Term premia and financial market indicators

One benefit of expressing the term premium as a single parameter  $\psi_t$  is that it can be used as the dependent variable in tests for the possible drivers of the term premium. I use a least squares regression of  $\psi_t$  on a range of money market indicators for New Zealand. This is by no means an exhaustive list, but it represents information about the money market that is freely available to market participants. (The Appendix contains descriptions of the variables and the data sources.) The regression equation is:

$$\psi_t = c + \gamma_1 \text{BIDCOVER}_t + \gamma_2 \text{OMO}_t + \gamma_3 \text{LEVEL}_t + \gamma_4 \text{SLOPE}_t + \gamma_5 \text{BILLVOL}_t + \gamma_6 \text{BONDVOL}_t + \varepsilon_t \quad (10)$$

The reasons for including these variables are as follows. The bid cover ratio for the weekly Treasury bill tender measures the demand for short-term government securities. During times of greater-than-usual uncertainty, or when banks' credit riskiness increases, the demand for these safer securities should increase relative to bank bills. The bid cover ratio should be positively correlated with the term premium.

The demand for cash in the RBNZ's daily liquidity operations increases when banks find it harder to raise cash in the bank bill market. This can happen when investors are close to or at their limits on how much bank-issued paper they can hold. Bank bill yields, particularly for maturities up to three months, tend to rise at these times. Therefore, the successful bid yields in open market operations are a proxy for one source of market illiquidity, and should be positively related to the term premium.

Both the level of interest rates and excess forward returns fell during the 1990's, so it is possible that the term premium also followed this trend. The slope of the yield curve depends on the stage in the monetary policy cycle, and there is some evidence that the premium increases during tightening phases and narrows during easing phases. This implies a positive relationship between the slope of the curve and the term premium, although the response is not necessarily symmetric.

Finally, the volatility of short-term rates is a proxy for interest rate risk. The historical volatility of bank bill yields is a backward-looking measure, while the volatility implied by option prices on short-term bonds is a forward-looking measure. Both measures should be positively related to the term premium.

Table 5 shows the results of the final estimates, with the insignificant variables removed from the regression. These indicators have fairly low explanatory power, accounting for less than a third of the variation in the term premium, and some of them have the incorrect sign.

The last two columns show the results using the three-month and twelve-month excess forward returns as the dependent variables. As

with the term premium, the indicators explain less than a third of the variation. This highlights the problem of having to choose the explanatory variables in a linear regression. By itself, this approach would have suggested that there is little evidence of time variance in the term premium. However, the Kalman filter estimates suggest that there is evidence of time variance, but the market variables used here may be poor proxies for what drives the premium.

**Table 5**  
**Term premia and money market indicators**

	(1)	(2)	(3)	(4)
Constant	0.047	0.076	0.546	2.838
T-bill bid cover ratio	-0.002	-	-	-
OMO spread	-	-	-1.628	-
Level of yield curve	-0.001	-0.008	-0.153	0.201
Slope of yield curve	-	-	-	1.184
Bank bill historical volatility	0.109	0.154	11.068	13.660
Short bond implied volatility	-	-	-	-0.283
Adjusted $R^2$	0.319	0.265	0.329	0.313

Key of dependent variables:

- (1)  $\psi_t$ , mean-reverting specification.
- (2)  $\psi_t$ , non-stationary specification.
- (3) Excess forward returns,  $j=3$ .
- (4) Excess forward returns,  $j=12$ .

Other tests, not shown here, included using the term premia shock  $v_t$  as the dependent variable, and using the first differences of the dependent and/or explanatory variables. The parameter estimates generally had the same signs, and were less significant. The only variable that was consistently significant was the volatility of bank bill rates. This suggests that an approach such as the ARCH-M specification (Engle et al, 1987) would go some way to explaining changes in the term premium.

## 6 Conclusion

There is strong evidence of a time-varying term premium on short-term interest rates in New Zealand, with somewhat weaker evidence for Australia. The premium appears to be best described as slowly mean-reverting, although the results are not overly sensitive to assumptions about the degree of mean reversion. The variation in the premium is both statistically and economically significant.

It is not clear whether the method of monetary policy implementation has any influence on the size or variability of the term premium. During the periods when the central banks used the cash rate as their main tool, the term premia estimates were relatively stable, so much so that the mean-reverting estimates usually converged to a constant value. However, the small sample size makes it difficult to distinguish between the impact of the implementation method and the impact of one-off economic shocks.

The outputs from the Kalman filter method could be used to examine the market reaction to economic and policy shocks. For example, the estimates presented here end a few months after the 11 September 2001 terrorist attacks – an event that could have led market participants to reassess many of the risk factors that drive the term premium. As more data becomes available, it will be possible to estimate how quickly the market recovered from this shock, or whether it had a permanent impact.

This analysis could be refined by developing a better measure of market interest rate expectations. One possible way to do this is to include an expectations generation process in the model. This would mean that more of the variation in excess forward returns would be attributed to changes in market forecasts, and in turn it may lead to smoother estimates of the term premia. Another way is to use surveys of economists' forecasts, which are useful if not totally representative of the whole market. The author has survey data for both countries since 2000, but a larger data set would be necessary for any comprehensive analysis.

## Data appendix

### A.1 Data for term premia estimates

The sample periods examined here are those for which a full data set is available. The data set begins in November 1993 for Australia and March 1995 for New Zealand, and both data sets end in February 2003. After allowing for the number of lags in the estimation, the model produces estimates of the term premium for the longest horizon up to February 2002.

“Market cash rate” is the weighted average of the inter-bank lending rates, which is published every business day by the respective central banks. The market rate can differ from the policy rate, but the experience of the overnight cash rate regime is that these deviations are not large or persistent (Lowe (1995), Brookes and Hampton (2000)).

“Bank bill rates” are quoted for monthly horizons up to six months. Each business day, the inter-bank price makers submit the rates at which they believe the market is trading these bills. The respective Financial Markets Associations (NZFMA and AFMA) compile the results and publish the trimmed mean rates, which are used to value contracts such as futures and FRAs at the expiry date. “FRA rates” are forward contracts on the 90-day bank bill, which are quoted for monthly horizons with settlement dates four to nine months ahead. The rates used here are end-of-day values taken from Bloomberg.

The bank bill and FRA rates are used to build a yield curve of spot rates with up to 12 months to maturity, by compounding the appropriate spot and forward rates:

$$r_{t,b} = \left[ \left( 1 + r_{t,a} * \frac{a}{36500} \right) * \left( 1 + f_{t,b} * \frac{b}{36500} \right) - 1 \right] * \frac{36500}{b-a} \quad (\text{A.1})$$

Where  $a$  and  $b$  are the number of days to maturity for the relevant spot rates, and  $b > a$ . Forward one-month rates, with settlement dates one to eleven months ahead, can be calculated by rearranging

equation (A.1). The spot and forward rates were converted to continuously compounding rates, and the excess forward returns were calculated by subtracting the spot rate from the forward rate.

Strictly speaking, all of the rates must have the same length of time from settlement to maturity. To compare the one-day spot rate with forward one-month rates, I treat the average daily cash rate over the relevant month as the “spot” rate.<sup>12</sup> For example, the “spot” rate at the end of January is the average of the cash rate during February. As a result, the one-month-ahead forward rate ( $j=1$ ) is actually the one-month bank bill rate at the end of January, which is the market forecast of the average cash rate during February.

### A.2 Data for key money market indicators

“Bid cover ratio” is the ratio of bids received to the amount offered for the three-month Treasury bill in weekly tenders held by the RBNZ on behalf of the Crown. The data is taken from the last weekly tender of each month. The T-bill market is very illiquid and the tender is the only time when T-bills are traded in any significant volumes. The data is published weekly on the RBNZ website [www.rbnz.govt.nz](http://www.rbnz.govt.nz).

“OMO spread” is the weighted average successful bid in the RBNZ’s daily open market operations, minus the cash rate. The rates used here are for reverse repurchase agreements, where the RBNZ lends cash in exchange for collateral (this has been the typical transaction over most of the sample period). The rates are the monthly averages of daily observations. The data is published daily on the RBNZ website.

“Level” is the one-month average of the market cash rate at the end of the month.

“Slope” is the difference between the 90-day bank bill rate and the cash rate at the end of the month.

<sup>12</sup> Another approach, used in Krippner (2002) and Gordon (2002), is to use daily observations and interpolate forward one-day rates from the monthly rates.



“Bank bill historical volatility” is the 60-day rolling standard deviation of daily changes in the three-month bank bill yield. “Short bond implied volatility” is the annual volatility implied by the price of a three-month option on a government bond, typically with 2-3 years to maturity. The rates are quoted by the Bank of New Zealand and published on the Reuters page BNZWSWOP. Implied volatilities for bank bills and FRAs are also available, but they are updated much less frequently.

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