Testing the Expectations Theory of the Term Structure for New Zealand*

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Abstract

This paper tests the rational expectations theory of the term structure using recent daily, weekly, and monthly observations on New Zealand interest rates. We find that for many maturities we cannot reject the expectations hypothesis using both short and long versions of the theory. These results are interpreted as further evidence that the failure of the expectations hypothesis in the United States is due to the specific interest rate smoothing behaviour of the Federal Reserve.

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1 Introduction

The rational expectations hypothesis of the term structure is one of the fundamental building blocks of financial and macroeconomic theory. It has important implications for predicting future movements in interest rates, interpreting monetary policy, and in building macroeconomic models. An empirical literature, almost as extensive as the theoretical literature which assumes it, tests the expectations hypothesis using data from the US, and fails to find support for it; surveys are provided in Melino (1988),

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Shiller (1990), Campbell and Shiller (1991), and Campbell (1995). The most damning results occur when using the long version of the expectations hypothesis, where the term spread is used to predict future movements in long-term interest rates. In this case, not only is the expectations hypothesis rejected, but the spread predicts the wrong direction of the subsequent movement in the long rate.

A number of explanations have been proposed to explain the failure of the expectations hypothesis; these include irrational expectations (Froot, 1989), measurement errors (Stambaugh, 1988), small sample bias (Bekaert et al., 1997 and Schotman, 1997) and the Fed's particular interest rate smoothing behaviour (Mankiw and Miron, 1986). A useful way to differentiate between this last explanation and the others is through the use of pre-Fed data in the US, as well as international data. One would expect problems caused by irrational expectations, measurement errors, and small sample bias to be more substantial in countries other than the US and for the pre-Fed data. The post-Fed US market is more liquid and developed than either the pre-Fed US market or the markets that exist in other countries. If these were the only reasons for the failure of the expectations hypothesis, the failure in these other samples would be expected to be even more dramatic. Alternatively, if the Fed's specific interest rate smoothing behaviour was the sole cause of the failure, the expectations hypothesis should be more successful prior to the founding of the Fed and where central banks conduct monetary policy differently.

Some evidence which supports this interest rate smoothing explanation is provided by Mankiw and Miron (1986), who study the behaviour of three- and six-month interest rates in the US over the pre-Fed period, as well as the period subsequent to the founding of the Federal Reserve System in 1913. They find that the short version of the expectations hypothesis is supported in the pre-Fed period, but not in the later period. Hsu and Kugler (1997) also find the short version of the expectations hypothesis performs well for the US when using very recent data. They take a one-month maturity for the short rate and a three-month maturity for the long rate and consider the period October 1987 to November 1995. They argue the success of the expectations hypothesis for this later period is related to a change in Fed policy towards using the spread as an indicator for US monetary policy. Gerlach and Smets (1997) also find considerable support for the short version of the expectations hypothesis for a number of OECD countries. In particular, Belgium, Denmark, France, Italy, the Netherlands, Norway, Spain, and Sweden all had slope coefficients on the term spread which averaged close to one for the three maturities considered.[1] They argue the success of the expectations hypothesis in some countries, rather than others, is due to the greater degree of interest rate predictability in these countries; this in turn could be because central banks in these countries attach greater operational importance to exchange rate targets.

In this paper, we provide new evidence on the expectations hypothesis using data from a country not previously considered in this literature.[2] Data for New Zealand is particularly interesting for several reasons. Firstly, since 1989 the Reserve Bank of New Zealand (RBNZ) has been targeting inflation according to a fairly strict rule which does not allow much scope for interest rate smoothing.[3] Secondly, unlike many of the countries considered by Gerlach and Smets (1997), the exchange rate in New Zealand is freely floating, with no central bank intervention in the foreign exchange market. Finally, over the period we study, the RBNZ did not directly target short-term interest rates through open market operations. Instead, it allowed short-term interest rates to move freely with market expectations of the rates that were needed to best meet the Bank's inflation target. It did this by automatically linking its discount rate and

[1] The expectations hypothesis implies a coefficient on the term spread of one.

[2] An exception is Margaritis (1994), who finds some support for the expectations hypothesis using 32 quarterly observations of the 30-day and 60-day bank bill yields from 1985:1 to 1992:4 to estimate an ARCH-M model of the term structure.

[3] This approach can be summarized as: (a) the government assigns an explicit inflation target; (b) the central bank is given independence to achieve this target; (c) if the target is not met, the government requires an explanation from the Bank and has the option of dismissing the Bank's governor.

settlement cash interest rate to market rates. Open market operations were only used to meet a constant target for overnight cash settlement balances.[4] Under this set-up, the market immediately moved interest rates to offset the implications of any shock on future inflation. With large shocks, interest rates could move large amounts on a daily basis, despite there being no formal policy changes. Thus, the interest rate smoothing behaviour of the Fed, which Mankiw and Miron (1986) claim induces a near random walk in short-term interest rates, is unlikely to be an important feature of the data in New Zealand.

We run OLS regressions to test both the short and long versions of the expectations hypothesis, using New Zealand data. We consider a range of maturities with daily, weekly, and monthly data since 1989. The evidence is generally supportive of the expectations hypothesis. The point estimates of the slope on the term spread are centered around one, with many estimates insignificantly different from one. In no cases are point estimates negative, even for the long version. Moreover, using the forecasting equation proposed by Mankiw and Miron (1986), we find that interest rates are considerably more predictable in New Zealand compared to the US. Thus, this paper can be seen as further evidence for the hypothesis that the failure of the expectations hypothesis in the US is due to the specific interest rate smoothing behaviour that the Fed adopts.

The rest of the paper proceeds as follows. Section 2 gives a more detailed review of the expectations hypothesis and the different explanations for its failure with US data, together with the most recent empirical findings regarding these different explanations. Our data, estimation procedure, and pre-testing are outlined in Section 3, while Section 4 presents our main results. Finally, Section 5 briefly concludes.

[4]This policy worked because the RBNZ stood ready to force rates to its desired level if the market-delivered rates were not acceptable; such actions were rarely needed. For a more detailed description of this implementation procedure and an explanation of how it works, see Guthrie and Wright (1998).

2 A Review of Theory and Evidence

The rational expectations theory of the term structure asserts that the return on an n-period bond should equal the average of the expected returns on m-period bonds over the life of the n-period bond (where $n = km$), plus a term premium. That is,

$$R_t^{(n)} = \frac{1}{k}\sum_{i=0}^{k-1} E_t R_{t+mi}^{(m)} + \theta_{n,m},$$

where $R_t^{(n)}$ denotes the n-period return at time t. With the assumption of rational expectations, two regression models are usually used to test the expectations hypothesis. One involves predicting the future change of short-term rates using the term spread (short version)

$$\frac{1}{k}\sum_{i=1}^{k-1}(R_{t+mi}^{(m)} - R_t^{(m)}) = \alpha_s + \beta_s(R_t^{(n)} - R_t^{(m)}) + \varepsilon_{s,t}, \tag{2.1}$$

where $\varepsilon_{s,t} = \frac{1}{k}\sum_{i=1}^{k-1}(R_{t+mi}^{(m)} - E_t R_{t+mi}^{(m)})$, while the other predicts long-term rates using the term spread (long version)

$$R_{t+m}^{(n-m)} - R_t^{(n)} = \alpha_l + \frac{\beta_l}{k-1}(R_t^{(n)} - R_t^{(m)}) + \varepsilon_{l,t}, \tag{2.2}$$

where $\varepsilon_{l,t} = R_{t+m}^{(n-m)} - E_t R_{t+m}^{(n-m)}$.

When interpreting the results of these two regression approaches, three interpretations of the expectations hypothesis are often considered. In the first case, known as the pure expectations hypothesis, the term premium is zero. This corresponds to the hypothesis that $\alpha = 0, \beta = 1$. The second case allows the term premium to differ from zero, which corresponds to the hypothesis $\beta = 1$; we refer to this as simply the expectations hypothesis. The final hypothesis is that movements in future rates (either short- or long-term rates) are in the same direction as suggested by the expectations hypothesis. That is, $\beta > 0$. Clearly this is a much weaker requirement.

Campbell and Shiller (1991) provide a comprehensive study of the expectations theory of the term structure for US data. Using monthly data from January 1952 to

February 1987, they examine all possible pairs of maturities in the range 1, 2, 3, 4, 6, and 9 months and 1, 2, 3, 4, 5, and 10 years. They find that the slope on the term spread between almost any two maturities gives the wrong direction of forecast for the long version (equation (2.2) above). The coefficients are significantly different from one at conventional significance levels. For the short version of the theory, equation (2.1) above, the results are more mixed. In this case, the expectations hypothesis is rejected when the long-rate is less than 3 or 4 years. For longer maturities of the long-rate, the expectations hypothesis cannot be rejected for this version of the theory. Mankiw and Miron (1986) examine earlier episodes for the US and find that, prior to the forming of the Fed in 1913, the short version of the expectations hypothesis works well with interest rates of three- and six-month maturities. Hsu and Kugler (1997) also find the short version of the expectations hypothesis performs well when using very recent data. They take a one-month maturity for the short rate and a three-month maturity for the long rate and consider the period October 1987 to November 1995. Considerable support for the short version of the theory is also found by Kugler (1990) in the case of Germany and Switzerland, and by Gerlach and Smets (1997) in the case of Belgium, Denmark, France, Italy, the Netherlands, Norway, Spain, and Sweden. However, little positive evidence for the long version of the expectations hypothesis exists.

How can the empirical failure of the expectations hypothesis for the US be explained and how is this consistent with the supportive evidence for the short version of the expectations hypothesis from certain periods and certain countries? One explanation for the failure is that investors are irrational. The usual test of the expectations hypothesis is in fact a joint test of two different hypotheses. One hypothesis is that investors are risk neutral and the other is that investors' expectations are rational. However, using survey data on interest rate expectations, Froot (1989) was able to directly test the rationality hypothesis. He finds that the failure of the expectations hypothesis at long maturities is due to the underreaction of expected future rates to changes in the short rate. According to this explanation, rational expectations but not the expectations theory is rejected. This raises the question as to how much we can trust survey data in this kind of study? It is worth noting that even Froot does not find evidence of irrationality for short maturities. A more fundamental problem with the irrational expectations hypothesis is the difficulty it faces in explaining why the expectations hypothesis is rejected so strongly in the period 1952–1987 for the US, but performs so much better in the pre-Fed and post-1987 periods, as well as for some smaller countries. A priori, one would expect irrational expectations to be more prevalent in the pre-Fed period and countries with less developed markets than post-war US. Moreover, there is no apparent reason for the rationality of expectations to have changed around 1987 in the US.

Measurement errors in long-term rates has been suggested by Stambaugh (1988) as another potential explanation for the failure of the expectations hypothesis. The regression (2.2) is sensitive to measurement errors in long-term interest rates since the long-term rate appears both in the regressor with a positive sign and in the dependent variable with a negative sign. Hence measurement errors reduce the value of the coefficient on the term spread, and could be responsible for the negative signs that arise in tests of the long version of the theory. Using instrumental variables to handle the measurement errors, Campbell and Shiller (1991) find that the spread still predicts the wrong direction in the change of long-term rates. Also using an instrumental variables approach to handle measurement errors, Hardouvelis (1994) finds a similar result for the US. Hardouvelis also considered the other six G7 countries and finds that the point estimates in the long version are now positive, although still well below one.[5] Given that measurement error would seem to be more important in earlier times, as well as

[5]Only the OLS point estimates for France and Italy are positive; these are still well below the value predicted by the expectations hypothesis. Point estimates from IV estimation are positive for all countries except the U.S., although still well below one. The standard errors of Hardouvelis' estimates are too large to render statistical testing informative.

in countries with less liquid markets, this explanation fails to explain the cross-country results discussed above.

Since the finite sample distribution of the estimator and the regression tests depend on the dynamic behaviour of short-term interest rates, small sample biases are examined in the literature. Bekaert et al. (1997) find that the persistence of short-term interest rates induces extreme bias and extreme dispersion into the small-sample distributions of the conventional test statistics usually employed. Schotman (1997) finds that when the short-term interest rate follow a process (say ARMA(1,1)) which is close to a random walk, the OLS slope coefficient can be badly biased in small samples. Instead of assuming short-term interest rates follow a random walk, Valkanov (1998) uses a local-to-unit process to model the short-term interest rate and correspondingly derives the alternative distributions for the regression tests. However, inferences based on the finite sample distributions of the specification tests statistics still provide a consistent rejection of the expectations hypothesis; see Bekaert et al. (1997) and Valkanov (1998). Moreover, the success of the expectations hypothesis for certain small sub-samples of the US data, as well as for some other countries, where sample sizes are less than half the size of the typical US studies, suggests this small sample bias is not driving the various results.

Fama (1984) and Mankiw and Miron (1986) show that a time-varying risk premium may destroy the predictive ability of the term spread. Under the maintained assumption of rational expectations, the probability limit of the estimated slope coefficient is less than one. The extent of downward bias depends on the relative importance of the variance in the risk premium relative to the variance of predictable changes in interest rates, and on the correlation of the risk premium with the term spread.[6] Thus, for a given process for the risk premium, less predictable interest rates imply a

[6] Tzavalis and Wickens (1997) explore the later possibility. One problem with assuming a risk premium which is correlated with the term spread is the difficulty in explaining why a positive correlation should exist.

greater downward bias in the slope coefficient. This gives rise to the notion of Mankiw and Miron, that when the Fed smoothes short-term rates it induces a random walk behaviour in these rates that makes interest rates less predictable and increases the downward bias in the slope parameter. Rudebusch (1995) formalizes the argument by showing that an empirical model of the Fed's interest rate targeting approach can give rise to the empirical results on the expectations hypothesis. He estimates a daily model of the Fed's interest rate targeting behaviour, which, accompanied by the maintained expectations hypothesis, explains the varying predictive ability of the yield curve and elucidates the link between Fed's policy and the term structure. Similar approaches have been adopted by Dotsey and Otrok (1995), McCallum (1994) and Hsu and Kugler (1997), although they use more ad hoc policy reaction functions to describe the central bank's choice of monetary policy.

Some further evidence which supports this interest rate smoothing explanation is provided by Mankiw and Miron (1986), who study the behaviour of three- and six-month interest rates in the US over the pre-Fed period, as well as the period subsequent to the founding of the Federal Reserve System in 1913. They find that the short version of the expectations hypothesis is supported in the pre-Fed period, but not in the later period. Hsu and Kugler (1997) also find the short version of the expectations hypothesis performs well for the US when using very recent data. They take a one-month maturity for the short rate and a three-month maturity for the long rate and consider the period October 1987 to November 1995. They argue the success of the expectations hypothesis for this later period is related to a change in Fed policy towards using the spread as an indicator for US monetary policy. Kugler (1990) finds support for the short version of the expectations theory using Euro DM and Euro franc interest rates, but not for the US. He attributes this result to the higher variability of expected interest rate changes in the German and Swiss case. This is consistent with the interest rate smoothing explanation, since both Germany and Switzerland target money growth rates rather than short term interest rates. Gerlach and Smets (1997) consider 17 countries and find that US short-term interest rates are the most difficult to predict and have the lowest slope parameters in a prediction regression. For the short version of the theory, they are unable to reject the expectations hypothesis in 35 cases from 51 regressions. They provide additional support for an interest rate smoothing explanation, by showing that countries which have more predictable interest rate variation also have slope parameters closer to unity. When interest rates contain more predictable variation, any time-varying term premium becomes less important, and the expectations hypothesis performs better.

3 Data and Methodology

We test the expectations hypothesis with recent New Zealand data from the RBNZ. Our sample period is January 1st, 1989 through October 31st, 1998. We choose to start our sample in 1989 since this is the year the Reserve Bank Act was introduced, under which the Bank was mandated to target "price stability." It is the year that is usually referred to as the beginning of the Bank's inflation targeting policy. For short maturities (1, 2, 3, 4, 5, and 6 months) we use bank bill rates rather than T-bill rates, since we cannot obtain daily data on T-bill rates before February 1997. There are three reasons why we think this is not a serious problem. Firstly, these are the rates the RBNZ itself refers to. Furthermore, in New Zealand the market for bank bills is much more liquid than that for T-bills. Finally, bank bill rates command only a small premium over T-bill rates and this premium is relatively stable over time, reflecting the consistently high credit rating of the banks that issue these bills.[7] Bank bills are discount instruments which pay no coupon. For rates with maturities of one year or more (1, 2, 5, and 10 years) government bond data is used. All rates are

[7] For instance, using available 1997 daily data from the RBNZ, the average premium on the one-month rate is 21.75 basis points (with a standard deviation of 1.82 basis points) and on the three-month rate is 21.80 basis points (with a standard deviation of 2.28 basis points).

based on continuously compounded yields, recorded at 11 am each day. Thus, there is a maximum of 2470 daily observations, 512 weekly observations, and 118 monthly observations; the four- and five-month rate data is only available since April 1991; the six-month rate data is only available since January 1991.

As is standard, we correct for correlations in the error terms of equations (2.1) and (2.2) using the method proposed by Newey and West (1987). Since we wish to consider daily, weekly, and monthly data (high frequency data enables us to expand our otherwise short sample), we need to calculate the order of the moving average process arising from the two equations.[8] It turns out the number of moving average terms in the error term depends on the frequency of the data, as well as the version of the hypothesis and the specific maturities considered. In the appendix, we calculate the appropriate order of the moving average process for each version of the theory. In (2.1), $\varepsilon_{s,t}$ follows an MA($n-m-1$) process, while in (2.2), $\varepsilon_{l,t}$ follows an MA($m-1$) process. It is important to note that m and n refer to the number of observations in the lifetime of the short- and long-term bonds, respectively. Thus as the frequency of observations increases, so does the number of moving average terms. We use the conventional Newey and West (1987) correction to ensure standard errors are adjusted for these autocorrelations.[9] For a reasonably long sample period it is common to have heteroskedasticity in both versions of the model tested. Thus the standard errors must also be corrected for heteroskedasticity; we use the method developed in White (1980). However, it should be noted that point estimates from our estimation do not depend on either type of correction above.

A number of problems emerge when rates with long maturities are used. Firstly,

[8] The literature predominantly uses monthly data. Two exceptions are Choi and Wohar (1991), who examine weekly, monthly, and quarterly data for a number of small sub-samples of US data, and Hsu and Kugler (1997), who examine daily, weekly, and monthly data for recent US data. However, in both cases, only the short version of the theory with one combination of maturities is tested.

[9] The performance of this correction will not be good when the degree of autocorrelation is large relative to the sample size; see Stock and Richardson (1989). This occurs for high n and low m in the short version of the theory and for high m in the long version of the theory.

the longer the maturities used, the more observations we have to throw away in the regressions. For instance, for the regression of 10-year rates against 5-year rates, we lose five years of data and only fifty observations are available for monthly data. Secondly, in estimating equation (2.2), $R_{t+m}^{(n-m)}$ is not always available. In these cases we use the standard approximation that $R_{t+m}^{(n-m)} = R_{t+m}^{(n)}$. When m is small relative to n, such an approximation is a good one. However, the validity of the approximation is questionable in the case of large m. Furthermore, because of the limited issuance of long-term bonds in New Zealand, the calculation of rates with long maturities by the RBNZ involves substantial measurement error.[10] For example, the government bonds on issue at the time of writing (January 1999) are February 2000, February 2001, March 2002, April 2004, November 2006, July 2009, February 2016. The five year bond rate is taken as the yield on the bond with the closest maturity, that being April 2004 in this case. However, the actual maturity of this bond is clearly greater than five years. Because of these approximation and measurement errors, when dealing with rates of long maturities one would expect some bias in the slope parameter estimates. A further problem may arise when using the term spread between rates less than one year with those one year and above. In this case, we are comparing yields on bank issued securities with those issued by the government. Despite the high quality of the bank issued securities, the risk premium on bank bills could still be correlated with changes in interest rates, in which case an additional bias will arise in the slope parameter estimates. A more serious concern is that long-term rates are based on continuously compounding yields, whereas the expectations hypothesis only strictly applies to zero-coupon bonds. A final problem may emerge if these various measurement problems induce non-stationarity in the variables used in equations (2.1) and (2.2).[11]

[10] Over the majority of the sample period the New Zealand government was running a budget surplus and paying back outstanding debt. Thus there was a thin market in government debt, relative to most other OECD countries, with only a limited number of maturities on offer.

[11] For instance, the term spread may be non-stationary when the maturity of the long rate is one year or greater and there is a term premium. To see this, take the case above of the bond issue which

In order to check for non-stationarity in the variables used in our regressions, we run augmented Dickey-Fuller unit root tests for all maturity combinations and frequencies considered. Although interest rates appeared to exhibit some downward trend over our sample, there is nothing in economic theory to suggest that nominal interest rates should exhibit a deterministic time trend. Therefore, we use a specification which allows for a constant term, but no trend. Tables 1, 2 and 3 record the results of the unit root tests corresponding to the different specifications of the expectations hypothesis considered in the paper. The main thing to note from these tests is that for all three frequencies and across a range of short rates, as the maturity of the long rate increases it becomes more difficult to reject non-stationarity of the series used in our regressions (since the ADF test statistic tends to increase as the maturity increases). As a consequence of this, as well as the measurement problems for rates with long maturities, the reader should interpret results which involve long maturities with considerable caution. In fact, the greater the maturity of the long rate, the less confidence we attach to the results presented.

4 Empirical Results

In estimating equations (2.1) and (2.2), we consider daily, weekly, and monthly frequencies for all combinations of maturities. Given the large number of results, Table 4 presents results where the maturity of the short rate is restricted to be one month. This allows us, in a snapshot, to compare the results across different versions of the theory, across different frequencies, and between the US and New Zealand.

The first thing to note from Table 4 is that point estimates across different frequencies are quite similar. This is despite the fact daily data has 21 times more observations

expires in April 2004. In April 1998 it is recorded as a five-year bond, although really its maturity is six years. By April 2000 its maturity is only four years, although it is still recorded as a five-year bonds. As the true maturity of the bond falls, the premium on the bond is likely to decline, so that the term spread used in equations (2.1) and (2.2) may not be stationary.

than monthly data. In itself this is suggestive that a small sample bias is not present for New Zealand. It also suggests that the choice of which frequency to use can be based on which frequency better allows us to draw definite inference, in the sense that we can end up with smaller standard errors and hence tighter confidence intervals. With this in mind, weekly data appears to be best, with substantially lower standard errors than either daily or monthly data.[12] For this reason, we concentrate on results for weekly data for New Zealand in the remainder of the paper.

We start by comparing results for the US (from Campbell and Shiller (1991)) with the results from New Zealand. When weekly data is used for New Zealand, the number of observations available is similar across the two countries. For the long version of the theory, the slope coefficients for the US regressions are all negative (except the one-month versus two-month regression where the coefficient is 0.002 with a standard error of 0.238), and all are significantly different from one at the 5% significant level. In contrast, for New Zealand the slope coefficients are all positive; seven out of nine are not significantly different from one, and five out of nine are significantly greater than zero. Similarly, for the short version of the theory, the New Zealand results provide a better match with the expectations hypothesis. Unlike the US, the coefficients are all significantly greater than zero. With the exception of the coefficient on the five-year long rate,[13] New Zealand point estimates are uniformly closer to unity. Despite this, the substantially lower standard errors for the weekly New Zealand regressions make it difficult to accept that the slope coefficients are equal to one. This is reinforced by the result that with both daily and monthly data, we cannot reject the expectations

[12]One advantage of using high frequency data is to extract more information from the same sample period. However, there may be a limit to how much more information can be extracted from a given sample as the frequency is increased. Also, the number of moving average terms needed to correct for autocorrelation in the error terms, increases in proportion to the extra observations gained by moving to a higher frequency. These two aspects may be the reason why a weekly frequency leads to lower standard errors than a daily frequency.

[13]Recall that our confidence in making inference falls as the maturity of the long rate increases, due to non-stationarity in the data and measurement problems.

hypothesis (slope coefficient equal to one) in seven out of eight cases. In contrast, with monthly data for the US, the expectations hypothesis is only accepted for maturities of three years and more. Another difference in the short version between the US results and those for New Zealand is that the U-shaped relationship for the US results (seen in Table 4 and discussed in Rudebusch (1995)), in which coefficients decrease and then increase in the maturity of the long rate, is not present for New Zealand.[14]

A more thorough examination of New Zealand results is presented in the following two tables, which present estimates of β for all combinations of maturities using weekly data. The results from the different maturity combinations provide additional support for the expectations hypothesis. For the long version, we cannot reject that $\beta = 1$ in 30 out of 34 regressions. For the short version, we cannot reject $\beta = 1$ in 11 out of 27 regressions. In more than two thirds of the cases where we do reject the expectations hypothesis, the estimate of β is greater than one. The median estimate of β over both versions is 1.058 (for the short version it is 1.091, and for the long version it is 0.989). As with Table 4, using data with daily or monthly frequencies makes it more difficult to reject the expectations hypothesis, but only because standard errors are higher.[15]

As we argued in the previous section, significant measurement error and non-stationarity data problems arise when using rates with long maturities. In fact, the extent of these problems increases with the maturity of the long rate. Consequently, the regression results for rates with maturities greater than one year should be treated with some caution, and those with five or ten year long rate maturities should probably

[14] If anything, the results for New Zealand appear to be increasing in the maturity of the long rate. Again, this might be related to the particular measurement and non-stationarity problems arising with New Zealand long maturity rates.

[15] We also found that the constant term in equations (2.1) and (2.2) is significantly different from zero in 51 out of 61 cases. In all but one of these cases the sign of the estimate was negative (the median estimate of α over both versions is -0.11), consistent with the idea that long maturity bonds have higher average returns than short maturity bonds (as would be predicted by the preferred habitat theory of Modigliani and Sutch, 1966), as well as with international evidence (see Gerlach and Smets, 1997). Because of this, we reject the *pure* expectations hypothesis in all but three cases for the long version and all but four cases for the short version.

be discounted altogether. To understand the impact of ignoring results with long maturities, a comparison of the results between the regressions involving only short-term rates and those involving the rates with maturities greater than 1 year is conducted. For the short version (Table 5), the median of the estimates of β in the regressions with only short maturity rates is 1.074. The median of the estimates of β in the regressions with long maturity rates is slightly higher at 1.286. However, the range of values with long maturities is clearly substantially larger, with the highest slope estimate 3.466 being for the five-year versus ten-year spread. For the long version (Table 6), the median of the estimates of β in the regressions with only short maturity rates is 0.992. By chance, the median of the estimates of β in the regressions with long maturity rates is the same value, 0.992. However, unlike the results from short maturity rates, the range of estimates is now very large (from 0.145 to 6.080), with slope estimates increasing in the maturities of the short and long rates used in the regressions.

We now explore whether the various theories which purport to explain the failure of the expectations hypothesis in the US can also explain its success using recent New Zealand data. We find that even with monthly data, where the sample size for the New Zealand regressions was at most 118 observations, the point estimates on the yield spread are centered around one. Thus the results from New Zealand are not consistent with the small sample bias explanation for the failure of the expectations hypothesis in the US. As we have detailed, measurement errors are likely to be even more of a problem with New Zealand data (at least when the maturity of long rates is greater than a year), so that they cannot explain why the expectations hypothesis performs better in New Zealand than the US. There is little reason to think that traders are more rational in New Zealand than the US, although we have no evidence on this; survey data, if available, could be used to examine this issue directly. Thus, out of the theories put forward in the literature, the most natural explanation for the differences between the US and New Zealand results is the very different nature of interest rate targeting across the two countries. In order to pursue this argument, we need to show that movements in interest rates are more predictable in New Zealand than in the US.

In the US, the lack of predictability in short-term interest rates is well documented (see Mankiw and Miron (1986) and Gerlach and Smets (1997), for instance). These authors argue that the lack of predictability in short-term interest rates is a symptom of the specific interest rate smoothing approach of the Fed. We follow the same approach as these authors and estimate a simple univariate forecasting equation,

$$\frac{1}{k}\sum_{i=0}^{k-1}(R_{t+i}^{(m)} - R_t^{(m)}) = \gamma_0 + \sum_{i=0}^{3}(\gamma_{1i}R_{t-i}^{(m)} + \gamma_{2i}(R_{t-i}^{(n)} - R_{t-i}^{(m)})) + \epsilon_t, \tag{4.3}$$

where $n = 3, 6, 12$ months and $m = 1, 3$ months. Like them, we use the R^2 from this regression as a natural measure of the predictability of changes in short rates. Table 7 shows that the R^2s in New Zealand are much larger than those in the US. This suggests Mankiw and Miron's interest rate smoothing hypothesis for the US does not apply to New Zealand.

5 Conclusion

The results of this paper suggest that the expectations hypothesis provides a good description of the term structure of interest rates in New Zealand during recent times. On average, the yield spread correctly predicts the movement in subsequent short and long rates over a range of maturities and frequencies. There was no systematic bias in this finding, although when longer maturity rates were used the range of results was substantially greater. This latter result seems to be linked to measurement problems for long maturity rates.

Unsurprisingly, there is clear evidence that long maturity bonds have higher average returns than short maturity bonds, as has been found in other countries. However, the success of the expectations hypothesis for New Zealand, especially using short maturity rates and when considering the long version of the hypothesis, stands in stark contrast to the results for the US. We argued the most likely reason for the difference in results across the two countries is the quite different monetary policy approaches adopted.

The paper raises a number of avenues for future research. One is to explore using the term spread in forecasting interest rates in New Zealand. The evidence presented in this paper suggests that the term spread provides useful information about future changes in interest rates in New Zealand. Another direction to explore, is why different monetary policy approaches lead to different degrees of interest rate predictability. While the defence of a currency under a speculative attack offers a natural explanation why some European countries have a high degree of interest rate predictability, it does not seem such a reasonable explanation for New Zealand. More generally, there is a need to have a theoretical understanding of the link between the operation of monetary policy and the predictability of interest rates. Rudebusch (1995), provides a start in this direction. He estimates a daily model of the Fed's interest rate targeting behaviour, which, accompanied by the maintained expectations hypothesis, explains the varying predictive ability of the yield curve and elucidates the link between Fed's policy and the term structure. However, future efforts should try to address cross country differences in results. Along these lines, an interesting possibility is that the degree of interest rate predictability might be linked to the degree of transparency regarding monetary policy intentions. This could explain the recent improvement in the performance of the expectations hypothesis for the US, as well as the success for New Zealand that we have documented here. If this argument holds more generally, the movement towards greater transparency in monetary policy across OECD countries implies that one would expect to find greater support for the expectations hypothesis using recent data for a number of different countries.

6 Appendix

Define m as the number of observations to maturity for the short rate, n as the number of observations to maturity for the long rate, and $k = n/m$ (which we assume to be an integer). For example, suppose the frequency is weekly, the short rate is the two-month interest rate and the long rate is the six-month interest rate. Then $m = 8$, $n = 24$, and $k = 3$. The error term in the short version, equation (2.1), is

$$\varepsilon_{s,t} = \frac{1}{k}\sum_{i=0}^{k-1}(R_{t+mi}^{(m)} - E_t R_{t+mi}^{(m)}),$$

and in the long version, equation (2.2), is

$$\varepsilon_{l,t} = R_{t+m}^{(n-m)} - E_t R_{t+m}^{(n-m)}.$$

First, consider the case of the short version. Then

$$E_t(\varepsilon_{s,t}\varepsilon_{s,t+j}) = \frac{1}{k^2}E_t\left(\left[\sum_{i=0}^{k-1}(R_{t+mi}^{(m)} - E_t R_{t+mi}^{(m)})\right]\left[\sum_{i=0}^{k-1}(R_{t+j+mi}^{(m)} - E_{t+j} R_{t+j+mi}^{(m)})\right]\right).$$

We want to find the value of x such that $E_t(\varepsilon_{s,t}\varepsilon_{s,t+x}) = 0$ for all $j > x$. Note that the terms $E_t R_{t+mi}^{(m)}$ are all constants, conditional on time t information. Thus, applying the law of iterated expectations to terms of the form $E_t E_{t+j} R_{t+j+mi}^{(m)}$ gives

$$E_t(\varepsilon_{s,t}\varepsilon_{s,t+j}) = \frac{1}{k^2}E_t\left(\left[\sum_{i=0}^{k-1} R_{t+mi}^{(m)}\right]\left[\sum_{i=0}^{k-1}(R_{t+j+mi}^{(m)} - E_{t+j} R_{t+j+mi}^{(m)})\right]\right).$$

Now when all the terms in the first square bracket are in the information set used to construct $E_{t+j} R_{t+j+mi}^{(m)}$, the term in the second square bracket will be orthogonal to the term in the first square bracket, and so this expression will equal zero. This will occur when $j \geq (k-1)m$, so that $x = (k-1)m - 1$. Thus the error term $\varepsilon_{s,t}$ will be moving average of order $(k-1)m - 1$. Substituting in $k = n/m$ implies $\varepsilon_{s,t}$ is MA$(n - m - 1)$. Using our example above with weekly observations, this implies the error term would be MA(15), as opposed to MA(3) if monthly data was used.

Now consider the case of the long version. This time

$$E_t(\varepsilon_{l,t}\varepsilon_{l,t+j}) = E_t\left([R_{t+m}^{(n-m)} - E_t R_{t+m}^{(n-m)})][R_{t+j+m}^{(n-m)} - E_{t+j} R_{t+j+m}^{(n-m)}]\right).$$

Using the same procedure as above, we find that

$$E_t(\varepsilon_{l,t}\varepsilon_{l,t+j}) = E_t\left([R_{t+m}^{(n-m)}][R_{t+j+m}^{(n-m)} - E_{t+j} R_{t+j+m}^{(n-m)}]\right).$$

When the term in the first square bracket is in the information set used to construct $E_{t+j}R_{t+j+m}^{(n-m)}$, the term in the second square bracket will be orthogonal to the term in the first square bracket, and so this expression will equal zero. This will occur when $j \geq m$, so that the error term $\varepsilon_{l,t}$ will be moving average of order $m-1$; that is, MA$(m-1)$. Note that for a given maturity of the short rate, m depends on the frequency of the data.

References

Bekaert, G., R. Hodrick, and D. Marshall. (1997). "On biases in tests of the expectations hypothesis of the term structure of interest rates," *Journal of Financial Economics*, 44, 309–48.

Campbell, J. (1995). "Some lessons from the yield curve," *Journal of Economic Perspectives*, 9(3), 129–52.

Campbell, J. and R. Shiller. (1991). "Yield spreads and interest rate movements: A bird's eye view," *Review of Economic Studies*, 58, 495–514.

Choi, S. and M. Wohar. (1991). "New evidence concerning the expectations theory for the short end of the maturity spectrum," *Journal of Financial Research*, 14, 83–92.

Dotsey, M. and C. Otrok. (1995). "The rational expectations hypothesis of the term structure, monetary policy, and time-varying term premia," *Federal Reserve Bank of Richmond, Economic Quarterly*, 81(1), 65–81.

Fama, E. (1984). "The information in the term structure," *Journal of Financial Economics*, 13, 509–28.

Froot, K. (1989). "New hope for the expectations hypothesis of the term structure of interest rates," *Journal of Finance*, 44, 283–305.

Gerlach, S. and F. Smets. (1997). "The term structure of euro-rates: Some evidence in support of the expectations hypothesis," *Journal of International Money and Finance*, 16, 305–21.

Guthrie, G. and J. Wright. (1998). "Market-implemented monetary policy with open mouth operations," mimeo, University of Auckland.

Hardouvelis, G. (1994). "The term structure spread and future changes in long and short rates in the G7 countries: Is there a puzzle?" *Journal of Monetary Economics*, 33, 255–83.

Hsu, C. and P. Kugler. (1997). "The revival of the expectations hypothesis of the US term structure of interest rates," *Economics Letters*, 55, 115–20.

Kugler, P. (1990). "The term structure of euro interest rates and rational expectations," *Journal of International Money and Finance*, 9, 234–44.

Margaritis, D. (1994) "Time-varying risk premia in the term structure of interest rates in New Zealand," *Applied Financial Economics*, 4, 111-20.

Mankiw, G. and J. Miron. (1986). "The changing behavior of the term structure of interest rates," *Quarterly Journal of Economics*, 101, 211–28.

Melino, A. (1988). "The term structure of interest rates: Evidence and theory," *Journal of Economic Surveys*, 2:4, 335–66.

McCallum, B. (1994). "Monetary policy and the term structure of interest rates," *National Bureau of Economic Research Working Paper*, No. 4938.

Modigliani, F. and R. Sutch. (1966). "Innovations in interest rate policy," *American Economic Review*, 56, 178–97.

Newey, W. and K. West. (1987). "A simple, positive semi-definite, heteroskedasticity and autocorrelation consistent covariance matrix," *Econometrica*, 55, 703–08.

Rudebusch, G. (1995). "Federal Reserve interest rate targeting, rational expectations, and the term structure," *Journal of Monetary Economics*, 35, 245–74.

Schotman, P. (1997). "Small sample properties of the regression test of the expectations model of the term structure," *Economics Letters*, 57, 129–34.

Shiller, R. (1990). "The term structure of interest rates," in Friedman, B. and F. Hahn (eds.), *The Handbook of Monetary Economics*, (Amsterdam: North-Holland).

Stambaugh, R. (1988). "The information in forward rates: Implications for models of the term structure," *Journal of Financial Economics*, 21, 41–70.

Stock, J. and M. Richardson. (1989). "Drawing inferences from statistics based on multi-year asset returns," *National Bureau of Economic Research Working Paper,* No. 3335.

Tzavalis, E. and M. Wickens. (1997). "Explaining the failures of the term spread models of the rational expectations hypothesis of the term structure," *Journal of Money, Credit and Banking*, 29, 364–80.

Valkanov, R. (1998). "The term structure with highly persistent interest rates," mimeo, Princeton University.

White, H. (1980). "A heteroskedasticity-consistent covariance matrix estimator and a direct test for heteroskedasticity," *Econometrica*, 48, 817–38.

Table 1: Unit root tests for data in all three frequencies

Version Frequency	Short Monthly	Long Monthly	Short Weekly	Long Weekly	Short Daily	Long Daily
$n = 2$	−3.31	−3.55	−4.82	−4.91	−6.47	−6.47
	−3.32	−3.32	−4.30	−4.30	−6.56	−6.56
$n = 3$	−2.92	−3.39	−4.22	−4.33	−5.66	−5.90
	−3.25	−3.27	−4.09	−4.17	−5.24	−5.46
$n = 4$	−3.40	−3.98	−4.07	−3.53	−4.35	−4.99
	−3.16	−3.13	−3.38	−3.46	−4.08	−4.16
$n = 5$	−3.22	−3.97	−3.46	−3.48	−3.76	−4.97
	−3.14	−3.15	−3.25	−3.33	−3.85	−3.93
$n = 6$	−2.92	−4.17	−3.25	−3.50	−3.40	−4.90
	−3.07	−3.12	−3.19	−3.26	−3.76	−3.76
$n = 12$	−2.64	−3.30	−2.42	−4.17	−2.50	−5.88
	−2.93	−3.10	−3.27	−3.50	−4.10	−4.33
$n = 24$	−1.86	−3.61	−1.49	−4.36	−1.48	−5.80
	−1.09	−2.90	−0.94	−2.75	−2.16	−3.33
$n = 60$	0.22	−4.92	−0.20	−5.39	−0.31	−5.91
	−1.08	−2.37	−1.49	−2.37	−1.84	−2.74

Note: The augmented Dickey-Fuller test statistic is computed as $\hat{\tau} = \hat{\beta}/ase(\hat{\beta})$ in the model $\Delta X_t = \alpha + \beta X_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta X_{t-j} + \varepsilon_t$, where X_t represents either the LHS variable or the RHS variable in the term structure regressions. The first row is for the LHS variable and the second row is for the RHS variable. The short maturity in each case is one month. The value of p is chosen by AIC. The 10% critical value is -2.57. The 5% critical value is -2.86.

Table 2: Unit root tests in the short version. Frequency: weekly

$n \backslash m$	1	2	3	4	5	6	12	24	60
2	−4.82	·	·	·	·	·	·	·	·
	−4.30								
3	−4.22	·	·	·	·	·	·	·	·
	−4.09								
4	−4.07	−3.00	·	·	·	·	·	·	·
	−3.38	−3.61							
5	−3.46	·	·	·	·	·	·	·	·
	−3.25								
6	−3.25	−3.01	−2.43	·	·	·	·	·	·
	−3.19	−3.31	−3.47						
12	−2.42	−2.40	−2.51	−2.09	·	−1.97	·	·	·
	−3.27	−3.58	−3.29	−2.59		−2.74			
24	−1.49	−1.38	−2.51	−1.92	·	−2.56	−2.05	·	·
	−0.94	−2.41	−2.18	−1.75		−1.81	−2.26		
60	−0.20	0.32	0.51	−1.67	−1.60	−2.71	−0.91	·	·
	−1.49	−1.19	−1.08	−2.01	−2.06	−1.85	−1.28		
120	·	·	·	·	·	·	·	·	−2.14
									−2.26

Note: The augmented Dickey-Fuller test statistic is computed as $\hat{\tau} = \hat{\beta}/ase(\hat{\beta})$ in the model $\Delta X_t = \alpha + \beta X_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta X_{t-j} + \varepsilon_t$, where X_t represents either the LHS variable or the RHS variable in the term structure regression (2.1) for n months versus m months. The first row is for the LHS variable and the second row is for the RHS variable. The value of p is chosen by AIC. The 10% critical value is −2.57. The 5% critical value is −2.86.

Table 3: Unit root tests in the long version. Frequency: weekly

$n\backslash m$	1	2	3	4	5	6	12	24	60
2	−4.91	·	·	·	·	·	·	·	·
	−4.30								
3	−4.33	·	·	·	·	·	·	·	·
	−4.17								
4	−3.53	−2.96	·	·	·	·	·	·	·
	−3.46	−3.61							
5	−3.48	·	·	·	·	·	·	·	·
	−3.33								
6	−3.50	−2.99	−2.37	·	·	·	·	·	·
	−3.26	−3.34	−3.47						
12	−4.17	−3.32	−2.98	·	·	−2.17	·	·	·
	−3.50	−3.90	−3.47			−2.73			
24	−4.36	−3.67	−3.36	−4.37	·	−2.78	−2.09	·	·
	−2.75	−2.87	−2.57	−1.80		−1.79	−2.26		
60	−5.39	−4.26	−3.47	−4.35	−4.45	−2.62	−2.38	·	·
	−2.37	−2.33	−2.07	−1.35	−1.32	−1.33	−1.72		
120	−5.22	−4.44	−3.61	−2.99	−3.38	−3.57	−2.37	−1.48	−1.91
	−2.33	−2.13	−1.85	−1.14	−1.16	−1.15	−1.63	−1.74	−2.23

Note: The augmented Dickey-Fuller test statistic is computed as $\hat{\tau} = \hat{\beta}/ase(\hat{\beta})$ in the model $\Delta X_t = \alpha + \beta X_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta X_{t-j} + \varepsilon_t$, where X_t represents either the LHS variable or the RHS variable in the term structure regression (2.2) for n months versus m months. The first row is for the LHS variable and the second row is for the RHS variable. The value of p is chosen by AIC. The 10% critical value is −2.57. The 5% critical value is −2.86.

Table 4: Regression results for U.S. and New Zealand. $m = 1$ month versus n months

Country	U.S.	U.S.	N.Z.	N.Z.	N.Z.	N.Z.	N.Z.	N.Z.
Version	Short	Long	Short	Long	Short	Long	Short	Long
Frequency	M	M	M	M	W	W	D	D
Sample	52:1	52:1	89:1	89:1	89:1	89:1	89:1	89:1
	87:2	87:2	98:10	98:10	98:10	98:10	98:10	98:10
$n = 2$	0.501†	0.002	0.833†‡	0.665‡	0.773†	0.547†	0.845†‡	0.690†‡
	(0.119)	(0.238)	(0.245)	(0.491)	(0.101)	(0.203)	(0.134)	(0.269)
$n = 3$	0.446†	−0.176	0.855†‡	0.475‡	0.780†	0.294	0.826†‡	0.431‡
	(0.190)	(0.362)	(0.240)	(0.510)	(0.077)	(0.228)	(0.137)	(0.316)
$n = 4$	0.436†	−0.437	1.187†‡	1.111‡	0.780†	0.814†‡	1.118†‡	0.968†‡
	(0.238)	(0.269)	(0.249)	(0.727)	(0.083)	(0.312)	(0.147)	(0.440)
$n = 5$	·	·	1.229†‡	1.174‡	1.074†‡	0.954†‡	1.238†‡	1.108†‡
			(0.267)	(0.812)	(0.079)	(0.350)	(0.174)	(0.525)
$n = 6$	0.237	−1.029	1.135†‡	1.179†	1.211†	1.033†‡	1.195†‡	1.202†‡
	(0.167)	(0.537)	(0.295)	(0.894)	(0.079)	(0.380)	(0.203)	(0.592)
$n = 9$	0.151	−1.219	·	·	·	·	·	·
	(0.165)	(0.598)						
$n = 12$	0.161	−1.381	1.021†‡	0.910‡	1.174†	1.076†‡	1.091†‡	1.160†‡
	(0.228)	(0.683)	(0.286)	(1.020)	(0.079)	(0.430)	(0.248)	(0.698)
$n = 24$	0.302	−1.815	1.308†‡	0.455‡	1.305†	0.794‡	1.303†‡	0.836‡
	(0.212)	(1.151)	(0.448)	(1.331)	(0.094)	(0.596)	(0.424)	(1.022)
$n = 36$	0.614†‡	−2.239	·	·	·	·	·	·
	(0.230)	(1.444)						
$n = 48$	0.873†‡	−2.665	·	·	·	·	·	·
	(0.291)	(1.634)						
$n = 60$	1.232†‡	−3.099	2.529†	0.579‡	2.477†	0.531‡	2.443†	0.485‡
	(0.182)	(1.749)	(0.183)	(2.097)	(0.075)	(0.968)	(0.122)	(1.701)
$n = 120$	1.157†‡	−5.024	·	0.167‡	·	0.145‡	·	0.106‡
	(0.094)	(2.316)		(3.155)		(1.492)	·	(2.640)

Note: † indicates that β is significantly greater than 0 at 5% level and ‡ indicates β is insignificantly different from 1 at 5% level.

Table 5: Regression coefficients β in the short version for NZ. Frequency: weekly

$n \backslash m$	1	2	3	4	5	6	12	24	60
2	.773†	·	·	·	·	·	·	·	·
	(.101)								
3	.780†	·	·	·	·	·	·	·	·
	(.077)								
4	.780†	.951†‡	·	·	·	·	·	·	·
	(.083)	(.133)							
5	1.074†‡	·	·	·	·	·	·	·	·
	(.079)								
6	1.211†	1.246†	1.120†‡	·	·	·	·	·	·
	(.079)	(.107)	(.153)						
12	1.174†‡	1.088†‡	1.102†‡	.994†‡	·	.871†‡	·	·	·
	(.079)	(.089)	(.096)	(.092)		(.010)			
24	1.305†	1.286†	.855†	.838†	·	.838†	.994†‡	·	·
	(0.094)	(.102)	(.061)	(.069)		(.082)	(.120)		
60	2.477†	2.552†	2.611†	1.091†‡	.991†‡	2.198†	2.580†	·	·
	(.075)	(.068)	(.071)	(.229)	(.236)	(.239)	(.077)		
120	·	·	·	·	·	·	·	·	3.466†
									(.177)

Note: † indicates that β is significantly greater than 0 at 5% level and ‡ indicates β is insignificantly different from 1 at 5% level.

Table 6: Regression coefficients β in the long version for NZ. Frequency: weekly

$n \backslash m$	1	2	3	4	5	6	12	24	60
2	$.547^{\dagger}$	.	.	.	.	.	.	.	.
	(.203)								
3	.294	.	.	.	.	.	.	.	.
	(.228)								
4	$.814^{\dagger\ddagger}$	$.902^{\dagger\ddagger}$	.	.	.	.	.	.	.
	(.312)	(.267)							
5	$.954^{\dagger\ddagger}$	.	.	.	.	.	.	.	.
	(.350)								
6	$1.03^{\dagger\ddagger}$	$1.161^{\dagger\ddagger}$	$1.241^{\dagger\ddagger}$	.	.	.	.	.	.
	(.380)	(.326)	(.305)						
12	$1.076^{\dagger\ddagger}$	$1.557^{\dagger\ddagger}$	$1.771^{\dagger}$	.	.	$.657^{\dagger\ddagger}$	.	.	.
	(.430)	(.318)	(.273)			(.198)			
24	$.794^{\ddagger}$	$1.180^{\dagger\ddagger}$	$1.323^{\dagger\ddagger}$	$1.393^{\dagger\ddagger}$	.	$1.138^{\dagger\ddagger}$	$.989^{\dagger\ddagger}$	.	.
	(.596)	(.432)	(.376)	(.361)		(.291)	(.240)		
60	$.531^{\ddagger}$	$.760^{\ddagger}$	$.809^{\ddagger}$	$1.195^{\ddagger}$	$.904^{\ddagger}$	$1.178^{\dagger\ddagger}$	$.970^{\dagger\ddagger}$	.	.
	(.968)	(1.705)	(.609)	(.641)	(.587)	(.543)	(.366)		
120	$.145^{\ddagger}$	$.551^{\ddagger}$	$.619^{\ddagger}$	$1.018^{\ddagger}$	$.652^{\ddagger}$	$1.058^{\ddagger}$	$1.013^{\ddagger}$	$4.299^{\dagger\ddagger}$	$6.080^{\dagger}$
	(1.492)	(1.076)	(.945)	(.934)	(.858)	(.797)	(.617)	(2.534)	(.328)

Note: † indicates that β is significantly greater than 0 at 5% level and ‡ indicates β is insignificantly different from 1 at 5% level.

Table 7: Predictability of changes in the short-term interest rate

	n	3 months	6 months	1 year	6 months
	m	1 month	1 month	1 month	3 months
	Frequency	Monthly	Monthly	Monthly	Monthly
NZ	89:1–98:10	0.15	0.41	0.37	0.32
US	64:6–93:12	0.08	0.11	0.15	·
US	59:1–79:2	·	·	·	0.03

Note: The U.S. results are from Gerlach and Smets (1997) and Mankiw and Miron (1986).