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Monetary policy implementation and uncovered interest parity: Empirical evidence from Oceania*

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Abstract

The close integration of Australian and New Zealand financial markets and the similarity of the monetary policy regimes provide the perfect backdrop for testing the empirical relevance of uncovered interest rate parity (UIP) in Oceania. We find that changes in the bilateral exchange rate have become more sensitive to the short-term interest differential over time. Most important, after the introduction of the Official Cash Rate regime in New Zealand, the responsiveness of the exchange rate has accelerated to such an extent that it is incompatible with UIP. Evidence on UIP over longer horizons is mixed with a 10-year horizon since 1990 providing the strongest support for the theory.

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1 Introduction

Uncovered interest rate parity (UIP) is a cornerstone condition in most open-economy macro models of the business cycle. In essence, it is an arbitrage condition that rules out excess profits in asset markets: the interest rate differential between the home and the foreign country equals the expected change in the exchange rate. Despite its pre-eminence in theoretical macroeconomic analysis, there is surprisingly little empirical evidence that UIP holds in practice. Surveys of the empirical literature in the 1980s and 1990s were unanimous in their rejection of the UIP hypothesis (Hodrick, 1987, Froot and Thaler, 1990, Lewis 1995, Engel 1996 to name but a few). Rather than finding the exchange rates of high-interest rate countries depreciating over relatively short investment horizons, most studies reported just the opposite: the exchange rates of high interest rate countries tended to appreciate. A recent comprehensive study by Burnside et al (2008) comes to the same conclusion.

Various explanations for the failure of UIP have been offered. One explanation focuses on the failure of two critical behavioural assumptions: that market participants are risk neutral and that they have rational expectations about future exchange rate changes. If market participants are not risk neutral, they will require a risk premium to hold foreign assets over domestic assets. Alternatively, if they systematically make errors when forming expectations of future exchange rates, potential risk-free profits from arbitrage will go unrealised.

Some commentators attribute the failure of the UIP hypothesis in practice to the conduct of monetary policy. McCallum (1994) shows that standard empirical tests of UIP produce negative results in models where the central bank uses the short-term interest rate as a policy instrument to respond to exchange rate changes. Building on this argument, Meredith and Chinn (2004) claim that the conduct of monetary policy makes short-term interest rate differentials highly volatile, thus contributing to the failure of UIP when tested over short horizons. They find much more support for UIP over 5- and 10-year horizons. Lothian and Wu's (2005) analysis of UIP over two centuries also backs up the theory, especially after they correct for the effect of regime changes on the formation of expectations.

More recently, the existence of carry trades between low- and high interest rate countries has been blamed for the failure of the UIP hypothesis. Indeed, as reported by the Economist (2007): "One obvious possibility is that the actions of carry traders are self-fulfilling: when they borrow the yen (the low interest rate currency) and buy the (US) dollar (the high interest rate currency), they drive the former down and the latter up. If other investors follow "momentum" strategies – jumping on the bandwagon of existing trends – this would tend to push up currencies with high interest rates."

Both the Australian Dollar and the New Zealand Dollar have been heavily involved in the carry trade. Both currencies have served as target currencies on account of the rather high interest rates in Oceania compared to those in countries like Japan or Switzerland whose currencies have served as funding currencies. If the carry trade is characterized by self-fulfilling tendencies as described above, then we would not be surprised to find that UIP does not hold between high and low interest rate countries. But the carry trade and its implications for UIP are not the focus of this paper. Rather, the objective of this study is to explore whether the arbitrage condition holds for two countries *between which* a large-scale carry trade is not operative.

We believe that the two largest countries in Oceania provide the perfect backdrop for an analysis of UIP because of their close economic integration. This is probably most obvious in the financial sector where Australian banks dominate. As a result, the average default risk on comparable bank-backed financial assets is similar in both countries. Other important factors are:

- the co-movement of interest rates over the business cycle in Oceania and the tendency of the New Zealand and Australian dollar to move in tandem vis-à-vis other major currencies.

- capital controls in Australia and New Zealand were largely abolished when the currencies moved to floating exchange rate regimes in 1983 and 1985, respectively.

- the central banks of both countries have similar monetary policy objectives. The primary focus of monetary policy is to keep inflation in check over the business cycle. Both central banks now implement monetary policy in a transparent way using almost identical operating procedures.

- until 2007 the lack of foreign exchange intervention on the part of the Reserve Bank of New Zealand (RBNZ) meant that the country's exchange rate represented one of the best examples of a 'clean' float.

- the geographic proximity and labour market flexibility stemming from the openness of both economies have aided the free flow of information and have provided the general perception of integration and stability.

The remainder of the paper proceeds as follows. The next section reviews basic issues with which researchers examining the UIP hypothesis have to grapple. Section 3 presents the empirical findings of our examination of UIP over both short and long horizons. Section 4 offers a more detailed explanation as to why the conduct of monetary policy might affect the arbitrage condition under CPI inflation targeting. Section 5 concludes the paper.

2 Theoretical Background

2.1 The Components of the UIP Hypothesis

Open economy macro models typically posit a relationship between nominal interest rate differentials and the nominal exchange rate. In these models domestic residents have a choice between holding domestic and foreign financial assets. Absent barriers to the free flow of capital, an arbitrage condition ensures that the expected return on both financial instruments is the same. A stochastic factor in the form of a risk premium is often added to this arbitrage condition. In log form, the relationship between the exchange rate, the interest rate differentials, and the risk premium is typically stated as follows:

$$\Delta s^{e}_{t,t+k} = i_{t,k} - i^{*}_{t,k} - \rho_t \quad (1)$$

$i_{t,k}$ = yield on domestic asset maturing in *k* periods

$i^{*}_{t,k}$ = yield on foreign asset maturing in *k* periods

$\Delta s^{e}_{t,t+k}$ = expected change in the exchange rate over the *k*-period interval

ρ_t = risk premium associated with holding foreign asset

The standard macro model assumes that expectations are formed rationally. Accordingly, the realized exchange rate in period *t*+*k* differs from its expectation by a white-noise error term:

$$s_{t+k} = s^{e}_{t,t+k} + \eta_{t+k} \quad (2)$$

s_{t+k} = exchange rate (measured in units of domestic currency per unit of foreign currency) in period *t*+*k*

$s^{e}_{t,t+k}$ = expected exchange rate for period *t*+*k* formed in period *t*

η_{t+k} = white-noise error that is orthogonal to all information dated *t*

Using (2), we can eliminate the unobserved expected exchange rate and restate (1) in terms of the observed change in the exchange rate from period *t* to period *t*+*k* and the difference between the yield on the domestic and the foreign asset, respectively. In addition, there is a composite error term that consists of the risk premium and the white noise error both of which are orthogonal to the interest rate differential:

$$\Delta s_{t,t+k} = i_{t,k} - i^{*}_{t,k} - \rho_t + \eta_{t+k} \tag{3}$$

Two variants of uncovered interest rate parity are typically tested in the empirical literature. The more narrow proposition assumes that agents form expectations rationally and that they are risk-neutral (no risk premium, i.e. $\rho_t = 0$). The less stringent proposition – the unbiasedness hypothesis – allows for the existence of a risk premium and non-rational expectations. The unbiasedness hypothesis merely requires that the risk premium and the forecast errors associated with the exchange rate be not systematically related to the interest rate differential.

2.2 Test of the Uncovered Interest Rate Parity Proposition

A simple linear regression equation forms the standard test of the UIP condition:[4]

$$s_{t+k} - s_t = \alpha + \beta(i_t - i^{*}_t) + \varepsilon_{t+k} \tag{4}$$

For UIP to hold, the probability limit of $\hat{\beta}$ must equal one. It is safe to say that most empirical studies emphatically reject this proposition. A general finding is that the estimated coefficient on the interest rate differential is significantly different from one – sometimes close to zero or even negative.

Fama (1984) and others[5] show that the existence of a variable risk premium can cause the coefficient on the interest rate differential to be different from its probability limit of unity. A few simple steps need to be taken to show this result. The OLS coefficient on the interest rate differential in (4) is given by:

$$\hat{\beta} = \frac{Cov(\Delta s_{t+k}, i_t - i^{*}_t)}{Var(i_t - i^{*}_t)} \tag{5}$$

Using (2) to replace the change in the exchange rate and (1) to substitute for interest rate differential allows us to restate the estimated coefficient in (5) as:

[4] For notational simplicity we drop the subscript *k* from both interest rates from here on. In addition, we ignore the subtle differences between the unbiasedness hypothesis and the UIP proposition.
[5] Hodrick and Srivastava (1984, 1986).

$$\hat{\beta} = \frac{Cov(\Delta s^e_{t+k} + \eta_{t+k}, \Delta s^e_{t+k} + \rho_t)}{Var(\Delta s^e_{t+k} + \rho_t)} \tag{6}$$

Assuming rational expectations allows us to simplify the above expression further:

$$\hat{\beta} = \frac{Var(\Delta s^e_{t+k}) + Cov(\Delta s^e_{t+k}, \rho_t)}{Var(\Delta s^e_{t+k}) + Var(\rho_t) + 2Cov(\Delta s^e_{t+k}, \rho_t)} \tag{7}$$

Inspecting (7), we note that the relationship between the risk premium and the expected change in the exchange rate plays a pivotal role in determining the sign and size of the estimated regression coefficient.

a. If $Cov(\Delta s^e_{t+k}, \rho_t) = Var(\rho_t) = 0$, (the probability limit of) $\hat{\beta}$ is unity.
b. If $Cov(\Delta s^e_{t+k}, \rho_t) = 0$, i.e. the risk premium is orthogonal to the expected change in the exchange rate, then $0 \leq \hat{\beta} < 1$ for $Var(\rho_t) > 0$. In this case $\hat{\beta}$ measures the variability of the expected change in the exchange rate relative to the sum of itself and the variance of the risk premium. The smaller is the variance of the expected change in the exchange rate relative to the variance of the risk premium, the smaller is the size of the estimated coefficient $\hat{\beta}$.
c. The estimated coefficient $\hat{\beta}$ can be negative. For this to occur, the covariance between the expected change in the exchange rate and the risk premium must be negative and greater in absolute size than the variance of the expected change in the exchange rate.
d. An interesting result emerges in the special case where $\hat{\beta} < 0.5$. The expression in (7) then reduces to $Var(\Delta s^e_{t+k}) < Var(\rho_t)$, i.e. the variance of the expected change in the exchange rate is strictly less than the variance of the risk premium.

3 Empirical Results

3.1 UIP over Short Horizons

This section presents the findings of our empirical investigation of the UIP condition between New Zealand and Australia from July 1986 to September 2008. We probe the validity of the UIP condition over a short horizon by initially examining the relationship between the differential on 90-day bank bill rates in New Zealand and Australia and observed exchange rate movements during the 90-day period (measured in calendar days). Following this, we test the same hypothesis using survey data on bilateral exchange rate expectations.

The economic landscape of New Zealand underwent dramatic changes in the 1980s. The decision to float the Kiwi Dollar in March 1985 was only the first of a series of important decisions that initiated a complete restructuring of the monetary policy framework. The complete overhaul of the monetary policy framework in New Zealand in the mid to late 1980s has been well-documented.[6] Arguably, the most important changes were the granting of independence to the Reserve Bank in 1989 and the formal announcement that the containment of inflation was the overriding concern of monetary policy.

The implementation of monetary policy in practice underwent dramatic changes in the late 1990s as well. On March 17th, 1999, the Reserve Bank introduced a new operating procedure that is centered on setting a short-term interest rate, the official cash rate (OCR). The Bank announces a change in the tune of monetary policy by altering the level of the OCR. Prior to the introduction of the OCR system, the Bank had employed an alternative operating procedure, one that featured a target for cash settlement balances (CSB). The quantity-based CSB framework proved rather unworkable. Its operation was marked by communication problems between the Reserve Bank and market participants and led to sizeable fluctuations in short-term interest rates as the financial markets were relied upon to play a pivotal role in setting short-term interest rates.[7]

Owing to the changes wrought upon the monetary policy framework in New Zealand over the 1986-2008 period, we subdivide the whole sample period into several sub-sample periods. They are:

July1st, 1986 – January 31st, 1990: period prior to Reserve Bank Act taking effect.

February 1st, 1990 – March 16th, 1999: operation of the cash settlement balances target framework.[8]

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[6] A descriptive analysis of the reforms and the circumstances which led to the adoption of a formal inflation target can be found in Reddell (1999), Sherwin (1999), and Guender and Oh (2006).

[7] Archer, Brookes, and Reddell (1999) describe the reason for the Reserve Bank's decision to switch from the CSB target to the OCR framework.

[8] For lack of a better term, we refer to this period as the cash settlement balances target period even though the Reserve Bank changed the target rather infrequently and often relied on "open-mouth operations" (Guthrie and Wright, (2000)) instead. For some time during

March 17th, 1999 – September 1st, 2008: operation of the official cash rate framework.

Another subsample period from February 1st, 1990 to September 1st, 2008 covers the period since the Reserve Bank Act went into effect. We add this subsample period to our analysis of UIP to determine whether the gain of independence by RBNZ had a material effect on the hypothesized link between changes in the exchange rate and the interest rate differential.

Table 1 provides summary information about the behaviour of the 90-day bank bill rates in New Zealand and Australia and changes in the bilateral exchange rate over the whole sample period and the subsample periods. Over the whole sample period the yield on short-term financial paper was higher on average and more variable in New Zealand than in Australia. In each subperiod, too, the mean return on 90-day bank bills was higher in New Zealand compared to Australia. Short-term interest rates in New Zealand were, however, less variable than in Australia during the Independence period (1990:Q1 – 2008:Q3) and also during the pre-OCR period (1990:Q1 – 1999:Q1). In New Zealand, short-term interest rates declined markedly from the first subsample period (1986:Q1 – 1989:Q4) to the third subsample period (1999:Q2 – 2008:Q3). Short-term interest rates also became more stable in New Zealand. Their variability during the third subsample period had reduced to one third of their variability in the first subsample period. A similar pattern exists for the behaviour of short-term interest rates in Australia except that short-term interest rates were more variable there in the decade spanning the 1990s than in the mid-to-late 1980s. Variability of changes in the bilateral exchange rate was most pronounced in the sub-sample period that followed the floating of the NZ Dollar. While the variability of short-term interest rates declined markedly in both countries during the OCR period compared to the pre-OCR period, the variability of changes in the nominal bilateral exchange rate actually increased by more than 40 percent.[9]

Figure 1 graphs the daily 90-day interest rate differential $(i_t^{NZ} - i_t^{AU})$ over the whole sample period. Notice that the volatility of the short-term interest

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this period the Reserve Bank also attempted to communicate to the public the stance of monetary policy through a monetary conditions index.

[9] The dramatic surge in carry trades that began in 2001 may have contributed somewhat to this increase in volatility. Being high-interest rate countries, both countries attracted sizeable inflows of short-term foreign investment. As a result, foreign exchange turnover in Australian and New Zealand dollars increased by 98 % and 152 % , respectively, over the 2001-2004 period (Galati and Melvin, (2004)). The Kiwi Dollar and the Australian Dollar remained carry trade target currencies also over the 2004-2007 period according to the 2007 BIS Triennial Survey of Foreign Exchange and Derivatives Market Activity.

rate differential between New Zealand and Australia during the OCR period was far lower than in the two preceding subsample periods. Inspection of column three in Table 1 reveals that the mean interest rate differential shrank from 197 basis points in the pre-independence period to 110 during the pre-OCR period and to 94 basis points during the OCR period. Of particular interest is the observation that the coefficient of variation (std. deviation/mean) decreases from 2.56 in the pre-OCR period to 0.66 in the OCR period.[10]

Table 2 presents the empirical results of estimating equation (4) for the entire sample period and for all subsample periods. Panels A – D show the regression results for daily, weekly, monthly, and quarterly data.[11] All regression results reported in these panels are based on actual observations of the nominal exchange rate. To ascertain whether survey data on exchange rate expectations tell a different story, we also test the UIP hypothesis on survey data. These findings are reported in panel E. In line with previous examinations of the UIP hypothesis, we focus primarily on the sign and size of the slope coefficient, i.e. the coefficient on the interest rate differential. Two simple hypothesis tests are carried out: $\beta = 0$ and $\beta = 1$.

For the case where actual exchange rate data enter the analysis, the estimation results over the whole period emphatically reject the UIP hypothesis. Our results for the 1986 – 2008 sample period are thus in broad agreement with the findings reported by previous studies. The coefficient on the interest rate differential is negative in all but one case and insignificantly different from zero for all data frequencies. The explanatory power of the

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[10] Figure 3 in the appendix superimposes the change in the quarterly exchange rate on the quarterly yield differential. It is apparent that the OCR period witnessed far smaller fluctuations in the yield differential than the other sub-periods while the volatility of exchange rate changes increased ever so slightly during the OCR period relative to the pre-OCR period. The summary measures of quarterly holding period returns and changes in the exchange rate are roughly the same as those reported for raw data in Table 1.

[11] We wish to state expressly that we present the empirical results based on daily and weekly data only for completeness sake. As such we do not wish to over-emphasize these results. Statistical inference based on daily and weekly data is hampered by the presence of ARCH and GARCH effects. The existence of overlapping observations also introduces a moving average error into the estimated regression equation. For daily and weekly observations this poses a serious problem as the standard errors are not consistent. Thus for daily and weekly data hypothesis tests of the statistical significance of regression coefficients are –strictly speaking – not valid. The problem is less severe for monthly observations ($k=3$ for monthly data on 90-day bank bill rates; the resulting MA error is of order $k-1 = 2$**).** All results reported were computed using OLS. Standard errors of the regression coefficient are based on the Newey-West procedure in EVIEWS with 2 or 3 truncation lags. The application of GMM where lags of the country-specific interest rates served as instruments yielded very similar results.

regression equation is virtually nil. Estimating the regression equation over the first two subsample periods, the pre-independence period and the pre-OCR period, respectively, yields results that are also at variance with the UIP hypothesis.

Estimating equation (4) over the last subsample period, the OCR period, produces dramatically different results. Inspection of the fourth row of panels A-D reveals that the estimated regression coefficient is positive and much greater than one for all data frequencies. For daily, weekly, and quarterly data β is significantly different from zero. Further, we find the hypothesis that the regression coefficient on the interest differential equals one can be rejected at the 5 percent level in daily and quarterly data. What is remarkable, to boot, is that the adjusted R^2 of all four regression equations estimated – while still low – is significantly higher for the OCR period than for the other two subsample periods. Thus, interest rate movements in New Zealand vis-à-vis Australia explain observed exchange rate movements to a much greater extent after the introduction of the new operating procedure in March 1999 than before.[12] Nevertheless, the conclusion emerges that observed movements in the NZ Dollar – Australian Dollar exchange rate were too excessive to be consistent with UIP in the OCR period.[13]

Panel E shows the regression results based upon quarterly survey expectations of the nominal NZ Dollar – Australian Dollar exchange rate. Survey expectations tell a different story. Over the whole sample period there is substantial support for UIP.[14] Not only is there a statistically significant positive relationship between the interest rate differential and expected changes in the exchange rate but the null hypothesis that the coefficient on the interest rate differential is unity cannot be rejected at

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[12] For the OCR period, averaging the daily data over the quarter reduces the size of the estimated coefficient on the interest rate differential. It also increases the standard errors, thereby producing lower t-values in Wald tests. The explanatory power of the regression based on averaged data is also lower compared to that of the regression based on last-observation-of-the-quarter data.

[13] Lothian and Wu (2005) propose an alternative way of testing the UIP hypothesis. It consists of pairing the yield on the foreign asset with the change in the exchange rate and then comparing the foreign return to the return on the domestic asset. Essentially, this approach requires, first, forming the sum of i_t^* and $s_{t+1} - s_t$ and, second, regressing this variable on i_t. The results remain largely the same with one exception. The null hypothesis that the coefficient equals unity cannot be rejected even though the estimated coefficient proves to be rather large at 3.133. Estimating the restricted version of the UIP test equation produces a much poorer fit as the adjusted R^2 drops markedly (0.032). It is debatable whether this approach improves upon the standard way of testing the UIP hypothesis. After all, it imposes the restriction that the coefficient on the foreign interest rate equals unity.

[14] Using survey data on expected exchange rates, King (1998) also finds support for the UIP hypothesis in Oceania over the 1987:Q2-1995:Q4 period.

conventional significance levels. The adjusted R^2 of 0.081 is also high relative to its counterpart in Panel D. Inspection of the results for each individual sub-sample period delivers a mixed picture. Despite its relative shortness, the first sub-sample period yields results that are broadly similar to those observed for the whole sample period. In marked contrast, there is no empirical evidence for UIP in the data covering the 1990:Q2 1999:Q1 subsample period. However, the introduction of the OCR framework again proves to have brought about a dramatic change in the relationship between the expected exchange rate change and the interest rate differential. For the OCR period, the results based on survey expectations match those based on actual observations of the exchange rate. The explanatory power of the regression surges to a relatively high adjusted R^2 of 0.313. But again, expected changes in the exchange rate react too sensitively to observed interest rate differentials during this period. The estimated coefficient on the interest rate differential shoots up to 8.168 which is far too large to be consistent with the UIP hypothesis.

To ascertain whether the change in the operating procedure in March 1999 coincided with the change in the relationship between the interest rate differential and the bilateral exchange rate, we varied the length of the OCR subsample period. To be specific, we both extended and shortened the OCR period by one or more quarters in an attempt to pinpoint the timing of the abrupt change in the sign and size of the regression coefficient. Panel A in Table *3* presents the results of this exercise for the case where the expected exchange rate is replaced with its observed value while Panel B does so for the case of survey-based expectations of the exchange rate.

According to Panel A, the size of the coefficient on the interest rate differential becomes statistically insignificant if the sample period for the regression estimated is lengthened by one quarter, i.e. by moving the beginning of the sample period from 1999:Q2 to 1999:Q1. Repeating this exercise for alternative starting dates of the last subsample periods generates similar results. In all cases, we find smaller estimates of the slope coefficient, all of which are statistically insignificant. In addition, the explanatory power of the regression is very low across the board prior to the introduction of the OCR framework in 1999:Q2. Beginning with 1999:Q2 and moving forward, we find a dramatically different picture. The size of the regression coefficient increases and remains statistically significant irrespective of the length of the last sub-sample period.[15] The other

[15] A predictive test for coefficient stability as suggested by Chow (1960) yields mixed results. The log likelihood ratio test rejects the null hypothesis of a stable coefficient on the interest rate differential (break point 1999:2) while the F-test does not. Adding to the regression equation a dummy variable which allows the coefficient to be different in size

noteworthy result is that the constant in the regression equation is statistically significant at the 10 or 5 per cent level for daily and quarterly frequencies during the OCR subsample period.[16]

For survey expectations, the increase in explanatory power of the regression equation begins prior to the introduction of the OCR regime. According to Panel B, the fourth quarter of 1998 marks the beginning of a continued jump in the adjusted R^2. The sudden change in the relationship between the interest rate differential and the expected exchange rate changes may simply reflect the fact that survey expectations of the bilateral exchange rate taken two quarters prior to the introduction of the OCR regime incorporated news about the impending change in the implementation of monetary policy in New Zealand.[17]

Taken altogether, we are led to the conclusion that uncovered interest rate parity is not a fitting description of the linkage between expected changes in the exchange rate and the short-term interest rate differential between New Zealand and Australia. UIP receives support from the data only if tested on survey expectations. Even then, support for UIP is concentrated in the early part of the whole sample period. The introduction of the OCR operating procedure caused a major change in the relationship between the interest rate differential and the expected change in the exchange rate. In more recent times, the Kiwi Dollar has indeed depreciated for positive interest rate differentials. Still, from the standpoint of UIP, movements in the observed exchange rate between two points in time have been too large for observed differences in short-term interest rates since the second quarter of 1999. We will return to this point in Section 4 of the paper in an attempt to provide a sensible explanation for this result. Before doing that, however, we examine the UIP hypothesis from a longer term perspective.

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during the OCR period leads to a positive but statistically insignificant coefficient on the interaction term. Other tests for structural stability (CUSUM, Chow breakpoint) produce no discernible evidence for instability.

[16] There is no serial correlation present in the residuals of the regression equations estimated for the OCR period.

[17] RBNZ formally announced its intention to switch to the OCR system on February 8th, 1999. Financial markets had not been given advance notice of the impending change in the operating procedure. However, it was no secret that RBNZ had contemplated making changes to the operating procedure as early as December 1996. These deliberations were put on hold when RBNZ introduced the monetary conditions index (MCI) as a communications device with the help of which RBNZ attempted to signal its intentions about the future course of monetary policy to financial markets and the public at large. Dissatisfaction with the performance of the MCI led to its eventual abandonment. For further details about the change in the operating procedure at RBNZ, see Archer et al (1999).

3.2 UIP over Longer Horizons

The examination of the UIP hypothesis over longer horizons follows largely the pattern of the previous section. There is one important exception, however. While the previous section considered interest differentials of bank bills, i.e. private short-term debt, this section considers interest differentials between longer-term government bonds. Two horizons are considered: a five-year and a ten-year horizon, respectively.[18] The conjecture is that UIP may hold over longer horizons because monetary policy is less likely to impact on long-term interest rates than short-term rates. In addition, long-term interest rates serve as a better gauge of inflationary expectations and risk than short rates.[19]

Table 4 and Table 5 present summary information about the behaviour of yields on 5-year and 10-year government debt. Apart from the 1990:Q1-2008:Q3 period in Australia, yields on 5-year government bonds were slightly more volatile than yields on 10-year bonds in both countries. New Zealand government bonds generated a higher mean return than Australian government bonds except during the pre-OCR period. The average interest rate differential on 5 year government bonds (0.35) was substantially higher over the whole sample period than the 10-year interest rate differential (0.11). During the OCR period, the volatility of the interest rate differentials was the same with the mean interest rate differential of the 5-year rate exceeding that of the 10-year rate by 11 basis points. As in the previous section, we take note of the steep decline of the coefficient of variation of the interest rate differential in the OCR period relative to the pre-OCR period. This normalized measure of dispersion falls from 5.47 to 0.48 for the 5-year interest rate differential and from 1.56 to 0.6 for the 10-year interest rate differential.

Turning attention to a comparison of 90-day interest rate differentials with 5- and 10-year differentials, we observe that interest rate differentials on short-term private paper were on average greater and more volatile than interest rate differentials on long-term public paper over the whole sample period and all sub-sample periods.

[18] The published long-term interest rates are only imperfect measures of the true yields on long-term government bonds. Ideally, one would use zero-coupon constant maturity interest rate series. Alexius (2001) proposes a method to remove the effects of coupon payments on bond prices.

[19] Mussa (1979), Alexius (2001), Razzak (2002), Chinn and Meredith (2004), and Lothian and Wu (2005) are proponents of this view. As pointed out in the introduction, risk is likely to be a minor factor in the current context.

A graph of the daily 5-year and 10-year interest rate differentials appears in Figure 2. It is apparent that both interest rate differentials were far more volatile prior to the introduction of the OCR regime. At the beginning of the whole sample period and at the end of the pre-OCR period both interest rate differentials were positive. In between they were often negative. It is also worth noting that the 5-year differential tended to be more volatile than the 10-year differential at the beginning of the sample period. In the latter stages of the pre-OCR period the 5-year differential also lay at times substantially above the 10-year differential.[20] During the OCR period, differences between the two interest rate differentials tended to be much smaller. Apart from two exceptions, the last four years of the sample period saw the 5-year differential lie above the 10-year differential in every month.

The empirical results of estimating equation (4) over a 5-year horizon appear in Table 6. Panel A reports the findings based on quarterly data while Panel B does so for annual data.[21] We also examine annual data to check the robustness of the quarterly results. Notice that the sample interval shrinks by 5 years.

The most striking result in Table 6 is that there is strong support for UIP in quarterly data over the whole sample period (1986:Q3-2003:Q4). The estimated coefficient on the interest rate differential is very close to one and highly statistically significant. Variations in the interest rate differential explain about 14 percent of the total variation in the first difference of the exchange rate.[22] However, further examination of the data reveals that the strong relationship does not carry over to the "Independence Period" (1990:Q2-2003:Q4) or to the OCR period. This suggests that the parity condition receives support from the data only in the early part of the sample period. Examination of annual data does not confirm the strong support the theory receives from quarterly data over the whole sample period.

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[20] Notice the surge of the interest rate differential during the Asian Currency Crisis (shaded area) when RBNZ initially tightened while the Reserve Bank of Australia eased the stance of monetary policy. At the time a monetary conditions index figured prominently in the monetary policy deliberations within RBNZ.

[21] The current change in the exchange rate could also be regressed on the lagged interest rate differential. Estimating this specification of the regression equation over the whole sample period yields results identical to those reported in Tables 6 and 7. Slight differences in the coefficient estimates emerge if the two specifications of the regression equation are estimated over sub-sample periods. However, these differences are not pronounced enough to warrant separate reporting. These results are available from the authors upon request.

[22] Estimating the same regression over the 1986:Q3-1999:Q1 period, i.e. the pre-OCR period, yields an adjusted R^2 of 0.241 and an estimated coefficient of 1.297 which is statistically significant at the 1 percent level.

For the 10-year horizon, the sample period extends from 1986:Q3 to 1998:Q4. Irrespective of how we specify the regression equation, the sample interval shrinks by 40 periods.[23] Table 7 records the estimated regression coefficients, the results of the hypotheses tests, and summary information about the fit of the estimated regression. Over the whole sample period the estimated regression coefficient is positive but statistically insignificant. The null hypothesis of $\beta=1$ can be safely rejected. Very different results are obtained for the shorter sub-sample period that begins in 1990:Q2, one quarter after the Reserve Bank of New Zealand gained its independence. For this sub-sample period the data provide strong support for UIP. The slope coefficient is estimated to be 0.759 and is significantly different from zero at the one percent level. The null hypothesis that the estimated coefficient equals one cannot be rejected at conventional significance levels. The explanatory power of the estimated regression rises threefold from 0.067 over the whole sample period to 0.211 over the 1990:Q2-1998:Q4 period. Using annual data, we find equally strong if not stronger evidence in support of UIP in the shorter sub-sample period compared to quarterly data. The estimated coefficient is very close to one and statistically significant. The adjusted R^2 surges to 0.390.[24]

Taken altogether, this section confirms to some extent the findings of Meredith and Chinn (2004), according to whom, evidence in favour of UIP is more likely to be found in the data when the theory is tested over long horizons. There is, however, no clear-cut evidence in Oceania that the parity condition holds systematically across time and over both 5-year and 10-year horizons. There is strong evidence in support of the UIP hypothesis over a 10-year horizon. But this support is limited to a sample interval that

[23] Alternatively, we could have specified the regression equation as:

$$s_t - s_{t-40} = \alpha + \beta(i_{t-40}^{NZ} - i_{t-40}^{AUS}) + \varepsilon_t$$

The sample period extends from 1996:3 to 2008:4 and includes the OCR period. The coefficient estimates, standard errors, test statistics and goodness of fit statistics are exactly the same for the whole sample period as reported in Table 7. For the OCR period (1999:2 - 2008:4) the regression results based on the alternative specification differ minutely from those reported in Table 7. For instance, the slope coefficient is now estimated to be 0.713 with standard error of 0.199.

[24] In Table 7 the constant in the regression is statistically significant over the whole sample period and the "Independence" period. This is indicative of the existence of a risk premium. Since the early 1990s the sovereign credit risk of both Australia (July, 1992) and New Zealand (April 1993) has been rated AAA by Standard and Poor's for long-term bonds issued in *local currency*. Fitch Ratings are identical though New Zealand's sovereign rating history with this agency extends back only to 2002. There are slight differences in the ratings of long-term debt issued in *foreign* currency with Australia enjoying AAA status since 2003. New Zealand has been rated AA+ since 1996. From 1996 to May 1999 New Zealand's credit rating for debt issued in foreign currency was slightly better than Australia's.

excludes the pre-Independence period. In marked contrast, there is strong evidence that UIP holds over a 5-year horizon in the early part of the sample period.

4 Interpreting the Short Horizon Results

In this section, we try to rationalize why the conduct of monetary policy may have played an instrumental role in the changing relationship between short-term interest rates and the nominal bilateral exchange rate in Oceania. At the outset we introduce a simple open-economy model. This model proves helpful in our attempt to provide a theoretical explanation for why monetary policy may create a systematic *positive* co-movement in *observed* exchange rate changes and the interest rate differential. After all, the robust positive linkage between the two variables has been observed only since the beginning of the OCR period when RBNZ formally switched to an interest-rate based operating procedure.

McCallum (1994) and Guender (2010) address the issue of the relationship between the instrument rule that the central bank follows to implement monetary policy and the behavior of the nominal exchange rate. In this section, we use a simplified open-economy model proposed by the latter author to explain why the change in the nominal exchange rate can be positively related to the lagged interest rate differential.

$$x_t = i_t - i_t^* \qquad (8)$$

$$x_t = E_t s_{t+1} - s_t + \rho_t \qquad (9)$$

$$x_t = \lambda_1 \pi_t^{CPI} + \lambda_2 x_{t-1} \qquad (10)$$

$$\pi_t = -\alpha x_t + \alpha(E_t \pi_{t+1} - E_t \pi_{t+1}^*) + u_t \qquad (11)$$

$$\pi_t^{CPI} = (1-\gamma)\pi_t + \gamma(\Delta s_t + \pi_t^*) \qquad (12)$$

The model consists of five equations. Variables marked by an asterisk denote the foreign counterpart of the domestic variable. Foreign variables are treated as exogenous random variables. The first equation is the definition of the policy instrument *(x_t)*. The policy instrument is defined as the difference between the domestic interest rate *(i_t)* and the foreign interest rate *(i_t^*)*. Adopting this convention simplifies the analysis and allows us to compare our results directly with those reported by McCallum. The second equation is the UIP condition in nominal terms. It allows for the existence of a risk premium *(ρ_t)*.

McCallum's model encompasses the first two equations and a version of equation (10). His version of the instrument rule features the change in the nominal exchange rate (Δs_t) instead of CPI inflation (π_t^{CPI}). He shows that interest rate smoothing by a central bank, i.e. $\lambda_2 \neq 0$, leads to a negative coefficient on the lagged interest rate differential in the reduced form equation for the change in the nominal exchange rate. This finding is consistent with the empirical evidence which typically rejects the validity of the UIP hypothesis. To be precise, the coefficient on the lagged interest rate in said equation equals the ratio of the policy parameters, i.e. $-\frac{\lambda_2}{\lambda_1}$. One would expect both policy parameters to be positive.

McCallum's policy target is clearly at odds with conventional practice. Most central banks have a target for the CPI inflation rate. They are less worried about changes in the nominal exchange rate. Indeed, the nominal exchange rate needs to be flexible as it acts as a shock absorber under inflation targeting. Hence, in line with equation (10), the policy instrument responds to deviations of the CPI inflation rate from target (which in the current set-up is assumed to be zero) and to its own lag.

To close the model, we add two equations, one for the rate of domestic inflation (π_t) and the other for the rate of CPI inflation. Equation (11) is a condensed equation of the rate of domestic inflation. It is obtained by combining three elements: an open economy IS equation, UIP in real terms, and a Phillips Curve. [25] The rate of domestic inflation is inversely related to the policy instrument, reacts positively to the difference between expected domestic inflation and expected inflation abroad in period *t+1,* and a stochastic disturbance. Equation (12) is the definition of CPI inflation.

The above model can be solved (via method of undetermined coefficients) for the change in the nominal exchange rate:

$$\Delta s_t = -\frac{\lambda_2}{\lambda_1 \gamma} x_{t-1} + \frac{(1+\lambda_1(1-\gamma)\alpha)}{\lambda_1 \gamma + \lambda_2} \rho_t - \frac{(1-\gamma)}{\gamma} u_t - \pi_t^* \qquad (13)$$

Equation (13) is evocative of the simple regression that is estimated in the standard empirical test of the UIP hypothesis employed in Section 2. The coefficient on the lagged interest rate differential depends on the policy

[25] We have used a simple Phillips Curve, i.e. one without forward-looking inflationary expectations. Likewise, the expected output gap next period has been dropped from the IS relation. These forward-looking expectations are not crucial to the results. Forward-looking expectations of the rate of domestic inflation do appear in equation (11), however, because in the IS relation the output gap depends inversely on the expected real rate of interest. The additive disturbance in equation (11) is a composite term consisting of random disturbances and exogenous variables that appear in the IS relation, the UIP relation, and the Phillips curve. For further details on the derivation of equation (11), see Guender (2010).

parameters λ_1 and λ_2 as well as the degree of openness γ. More precisely, the response of the change in the nominal exchange rate to the lagged interest rate differential depends on the ratio of the interest rate smoothing parameter to the parameter measuring the central bank's response to deviations of the target variable from its target level (as in McCallum's set up). The degree of openness also matters now as the target variable in the policy rule is not the change in the nominal exchange rate but the CPI inflation rate. The degree of openness determines the extent to which CPI inflation changes in response to a change in the nominal exchange rate or foreign inflation. As long as both policy parameters are positive, the coefficient on the lagged interest rate differential is negative

A simple change to the instrument rule overturns this result, however. Suppose that the central bank follows a more restricted version of the instrument rule (10):

$$x_t = \lambda_1 \pi_t^{CPI} + (1 - \lambda_1) x_{t-1} \qquad \lambda_1 > 0 \qquad (14)$$

According to this rule, the central bank has only one degree of freedom. It sets only one policy parameter with $\lambda_2 = 1 - \lambda_1$. Substituting for λ_2 in equation (13) changes the coefficient on the lagged interest rate differential in the following way:

$$-\frac{\lambda_2}{\lambda_1 \gamma} = -\frac{(1-\lambda_1)}{\lambda_1 \gamma} = -\frac{1}{\lambda_1 \gamma} + \frac{1}{\gamma} \qquad (15)$$

Here we see that the coefficient on the lagged interest rate differential in equation (13) need not be negative. Indeed, the sign of the coefficient depends on the size of the policy parameter λ_1. The more aggressively the central bank responds to CPI inflation deviations from target, the more likely it is that the coefficient in question becomes positive. Table 8 shows that for plausible values of the policy parameter λ_1 and the degree of openness (γ=0.4), the coefficient on the lagged interest rate differential takes on values ranging from a low of zero to a high around 2. Extreme reactions to CPI inflation deviations from target and a lower degree of openness such as γ=0.15 lead to an even larger coefficient on the lagged interest differential.

Thus, the large positive coefficient on the interest rate differential in the test equation during the OCR period may be a direct consequence of the way RBNZ implements monetary policy. At the very least, the positive coefficient is consistent with a scenario where the central bank:

1. follows an instrument rule which responds emphatically to deviations of CPI inflation from target; and

2. does not choose the response parameter on deviations from inflation from target independently of the smoothing parameter.

5 Conclusion

This paper examines the uncovered rate interest parity hypothesis in the two largest countries of Oceania over the 1986-2008 period. The empirical validity of this arbitrage condition is investigated against the background of a changing monetary policy framework in New Zealand. Three time horizons are analysed: a short time horizon covering 90-days and two longer time horizons covering 5 years and 10 years, respectively.

Our empirical examination of the UIP hypothesis detects a fundamental shift in the behaviour of observed exchange rate changes to the interest rate differential at the time the Reserve Bank of New Zealand adopted the OCR operating procedure in March 1999. Since then, exchange rate changes and the interest rate differential have been positively correlated to a significant degree. The introduction of a more transparent and efficient operating procedure whereby the policy instrument reacts systematically and swiftly to observed deviations of CPI inflation from target may have brought about this intriguing result. Naturally, other factors such as increased monetary policy coordination and the convergence of monetary policy objectives in the two countries may have been just as important.

Nevertheless, even with the adoption of the new monetary framework in New Zealand, UIP does not hold over a short horizon. The New Zealand Dollar is prone to depreciate excessively if the New Zealand 90-day bank bill rate exceeds its Australian counterpart.

Over longer horizons, the results are mixed. UIP cannot be rejected conclusively. Indeed we find some evidence in support of UIP over both 5-year and 10-year horizons. The problem with the evidence is that it is diffuse. UIP holds over a 10-year horizon but not over a 5-year horizon since the Reserve Bank became independent. There is some evidence in quarterly data that UIP holds even over a 5-year horizon but support for it is concentrated in the first five years of the sample period.

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Appendix

Figure 1
90-day bank bill interest rate differential

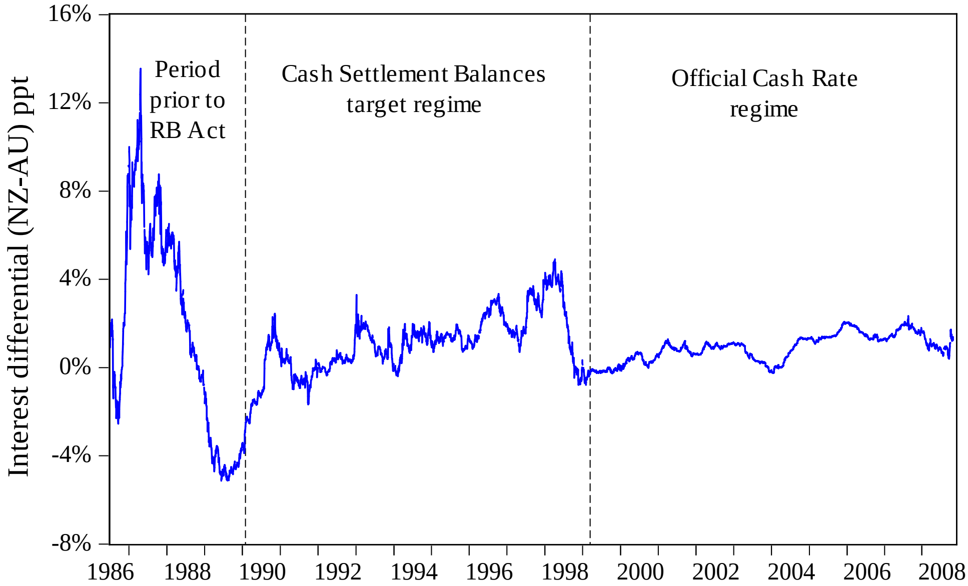


Figure 2
5- and 10-year interest rate differentials

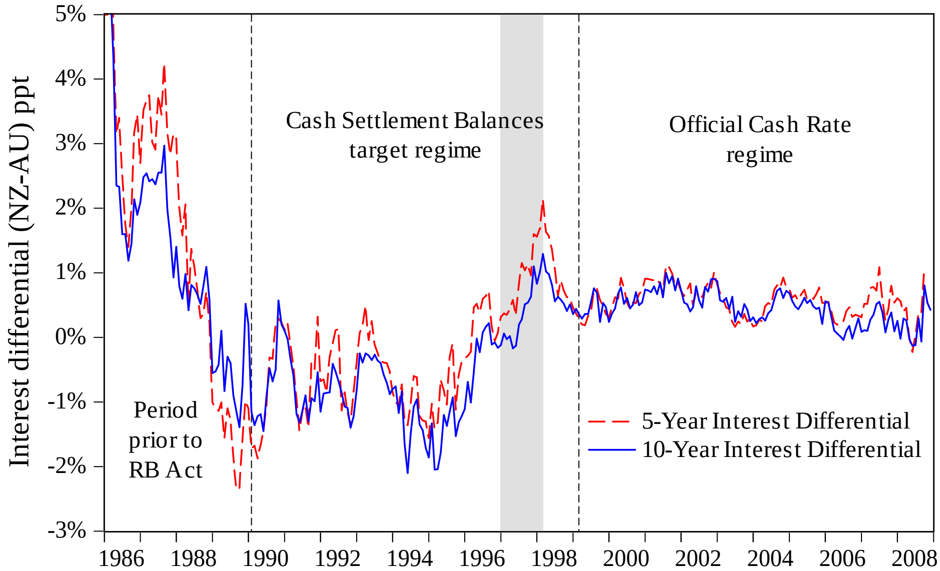


Table 1
Summary information for 90-day bank bills

	NZ90Day i_t^{90-day}	AUS90Day $i_t^{*90-day}$	$i_t^{90-day} - i_t^{*90-day}$	ΔEXCH Rate* (Δs_{t+1})
	Whole Sample Period (1986Q1 - 2008Q3)			
Mean	9.21	8.03	1.18	-0.0009
Std. Deviation	4.44	4.04	2.31	0.044
	Pre-Independence Period (1986Q1 - 1989Q4)			
Mean	17.10	15.13	1.97	0.0013
Std. Deviation	3.95	2.49	5.04	0.053
	Pre-OCR Period (1990Q1 – 1999Q1)			
Mean	8.43	7.33	1.10	-0.0015
Std. Deviation	2.46	2.91	1.42	0.032
	OCR Period (1999Q2 – 2008Q3)			
Mean	6.65	5.71	0.94	-0.0013
Std. Deviation	1.22	0.89	0.62	0.046
	Independence Period (1990Q1 2008Q3)			
Mean	7.53	6.51	1.02	-0.0014
Std. Deviation	2.12	2.27	1.08	0.043

Notes:

Table entries are based on quarterly data (last observation of the quarter). The NZ$/AUS$ exchange rate is expressed as units of NZ Dollars per Australian Dollar. Interest rates are nominal yields on 90-day bank bill rates.

*The sample period for exchange rate changes begins in 1986:3

Table 2
90-day bank bills

Panel A - Daily frequency

	α	β	Reject H_0: β=0	Reject H_0: β=1	$AdjR^2$	N
01/07/86-04/08/08	-0.000 (0.002)	-0.150 (0.271)	-	***	0	5765
01/07/86-31/01/90	0.005 (0.004)	-0.224 (0.272)	-	***	0.003	937
01/02/90-16/03/99	0.000 (0.002)	-1.006 (0.456)	**	***	0.012	2379
17/03/99-04/08/08	-0.009 (0.004)**	4.045 (1.383)	***	**	0.028	2449
01/02/90-04/08/08	-0.001 (0.002)	-0.074 (0.476)	-	**	0	4828

Panel B - Weekly frequency (Friday)

	α	β	Reject H_0: β=0	Reject H_0: β=1	$AdjR^2$	N
04/07/86-01/08/08	0.000 (0.003)	-0.129 (0.452)	-	**	-0.001	1153
04/07/86-26/01/90	0.004 (0.007)	-0.196 (0.464)	-	**	-0.003	187
02/02/90-12/03/99	0.001 (0.003)	-1.078 (0.766)	-	***	0.012	476
19/03/99-01/08/08	-0.011 (0.007)	4.853 (2.611)	*	-	0.034	490
02/02/90-01/08/08	-0.001 (0.003)	-0.085 (0.828)	-	-	-0.001	966

Panel C - Monthly frequency (last observation)

	α	β	Reject H_0: β=0	Reject H_0: β=1	$AdjR^2$	N
1986m07-2008m09	-0.001 (0.003)	0.003 (0.652)	-	-	-0.004	267
1986m07-1990m01	0.004 (0.012)	-0.042 (0.738)	-	-	-0.024	43
1990m02-1999m03	0.001 (0.005)	-1.118 (1.114)	-	*	0.005	110
1999m04-2008m09	-0.013 (0.009)	5.414 (3.629)	-	-	0.036	114
1990m02-2008m09	-0.001 (0.004)	-0.035 (1.101)	-	-	-0.005	224

Panel D - Quarterly frequency (last observation)

	α	β	Reject H_0: β=0	Reject H_0: β=1	$AdjR^2$	N
1986Q03-2008Q03	0.000 (0.003)	-0.216 (0.629)	-	*	-0.01	89
1986Q03-1990Q01	0.005 (0.012)	-0.445 (0.756)	-	*	-0.06	15
1990Q02-1999Q01	-0.000 (0.006)	-0.702 (1.642)	-	-	-0.024	36
1999Q02-2008Q03	-0.018 (0.009)**	7.238 (3.094)	**	**	0.054	38
1990Q02-2008Q03	-0.003 (0.005)	0.688 (1.323)	-	-	-0.011	74

Table 2 (cont.)
90-day bank bills

Panel E - Quarterly frequency **(Survey Expectations)**

	α	β	Reject H_0: β=0	Reject H_0: β=1	AdjR^2	N
1987Q03-2008Q03	**0.003 (0.003)**	**1.530 (0.370)**	*******	**-**	**0.081**	**86**
1987Q03-1990Q01	**0.009 (0.009)**	**1.709 (0.532)**	*******	**-**	**0.234**	**12**
1990Q02-1999Q01	-0.001 (0.004)	0.163 (0.740)	-	-	-0.029	36
1999Q02-2008Q03	-0.007 (0.005)	8.168 (1.227)	***	***	0.313	38
1990Q02-2008Q03	0.002 (0.003)	1.284 (0.915)	-	-	0.011	74

Notes:

1. Newey-West standard errors in parentheses. * (**) [***] indicates significance at the 10 (5) [1] per cent level. All daily or weekly dates are arranged in the following order: dd/mo/yr.

   Equation estimated:

$$s_{t+1} - s_t = \alpha + \beta(i_t^{NZ} - i_t^{AU}) + \epsilon_{t+1}$$

$$s_{t+1}^e - s_t = \alpha + \beta(i_t^{NZ} - i_t^{AU}) + v_t \quad (Panel\ E)$$

2. The survey expectation of the exchange rate is collected by the Reserve Bank and published on its website. The expectation is defined as the exchange rate of the NZ/AUS Dollar that survey participants expect to prevail at the end of the next quarter.
3. Simple compounding is used to convert annual yields to quarterly, monthly, and daily yields. For instance, the quarterly yield is computed as $i = \log(1 + x/4)$ where x represents the annual yield in decimal form.
4. The standard tests for non-stationarity were carried out prior to estimation. Over the whole sample period both variables - changes in the bilateral nominal exchange rate and the nominal interest rate differential - were found to be stationary.

Table 3
UIP and the adoption of the OCR regime in New Zealand: 90-day bank bill rates

Panel A – Actual Exchange Rate (Quarterly frequency)

	α	β	Reject H_0: β=0	Reject H_0: β=1	Adj R^2	N
1996Q01-2008Q03	-0.006 (0.007)	2.124 (1.510)	-	-	0.003	51
1997Q01-2008Q03	-0.007 (0.007)	2.731 (1.637)	-	-	0.011	47
1998Q01-2008Q03	-0.005 (0.007)	1.839 (2.119)	-	-	-0.012	43
1998Q04-2008Q03	-0.010 (0.011)	4.746 (3.702)	-	-	0.011	40
1999Q01-2008Q03	-0.012 (0.011)	5.276 (3.664)	-	-	0.017	39
1999Q02-2008Q03	**-0.018 (0.009)****	**7.238 (0.825)**	**	*	**0.054**	**38**
1999Q03-2008Q03	-0.022 (0.009)**	8.411 (3.039)	***	**	0.073	37
1999Q04-2008Q03	-0.024 (0.010)**	9.322 (3.190)	**	**	0.081	36
2000Q01-2008Q03	-0.024 (0.011)**	9.108 (3.535)	**	**	0.068	35
2001Q01-2008Q03	-0.027 (0.012)**	9.571 (3.14)	**	**	0.073	31

Panel B – Survey Expectations of Exchange Rate (Quarterly frequency)

	α	β	Reject H_0: β=0	Reject H_0: β=1	Adj R^2	N
1996Q01-2008Q03	0.007 (0.004)*	1.211 (1.292)	-	-	0.003	51
1997Q01-2008Q03	0.006 (0.005)	2.034 (1.724)	-	-	0.040	47
1998Q01-2008Q03	0.005 (0.005)	2.270 (2.332)	-	-	0.041	43
1998Q04-2008Q03	-0.005 (0.004)	4.561 (1.132)	***	***	0.299	40
1999Q01-2008Q03	-0.006 (0.004)	7.782 (1.166)	***	***	0.304	39
1999Q02-2008Q03	**-0.006 (0.005)**	**8.168 (1.227)**	***	***	**0.313**	**38**
1999Q03-2008Q03	-0.007 (0.005)	8.315 (1.370)	***	***	0.300	37
1999Q04-2008Q03	-0.007 (0.006)	8.212 (1.626)	***	***	0.268	36
2000Q01-2008Q03	-0.009 (0.006)	9..066 (1.681)	***	***	0.305	35
2001Q01-2008Q03	-0.007 (0.006)	8.314 (1.648)	***	***	0.240	31

Notes: Newey-West standard errors in parentheses). * (**) [***] indicates significance at the 10 (5) [1] per cent level.

Equation estimated:

$$s_{t+1} - s_t = \alpha + \beta(i_t^{NZ} - i_t^{AU}) + \epsilon_{t+1}$$

Table 4
Summary information for 5-year government bonds (nominal)

	NZ5Yr (i_t^5)	AUS5Yr (i_t^{*5})	$i_t^5 - i_t^{*5}$
	Whole Sample Period (1986Q3 - 2008Q3)		
Mean	8.29	7.94	0.35
Std. Deviation	3.17	3.08	1.13
	Pre-Independence Period (1986Q3 - 1989Q4)		
Mean	14.43	13.24	1.19
Std. Deviation	1.84	1.04	2.10
	Pre-OCR Period (1990Q1 – 1999Q1)		
Mean	8.12	8.29	-0.17
Std. Deviation	1.93	2.43	0.93
	OCR Period (1999Q2 – 2008Q3)		
Mean	6.21	5.65	0.56
Std. Deviation	0.51	0.54	0.27
	Independence Period (1990Q1 2008Q3)		
Mean	7.15	6.95	0.20
Std. Deviation	1.69	2.19	0.77

Table 5
Summary information for 10-year government bonds (nominal)

	NZ10Yr (i_t^{10})	AUS10Yr (i_t^{*10})	$i_t^{10} - i_t^{*10}$
	Whole Sample Period (1986Q3 - 2008Q3)		
Mean	8.25	8.14	0.11
Std. Deviation	2.99	2.97	0.92
	Pre-Independence Period (1986Q3 - 1989Q4)		
Mean	14.02	13.09	0.93
Std. Deviation	1.4	0.73	1.18
	Pre-OCR Period (1990Q1 – 1999Q1)		
Mean	8.15	8.69	-0.54
Std. Deviation	1.90	2.33	0.84
	OCR Period (1999Q2 – 2008Q3)		
Mean	6.22	5.77	0.45
Std. Deviation	0.47	0.46	0.27
	Independence Period (1990Q1 2008Q3)		
Mean	7.17	7.21	-.04
Std. Deviation	1.67	2.21	0.80

Table 6
5-year government bonds

Panel A – Quarterly frequency

	α	β	Reject H_0: β=0	Reject H_0: β=1	Adj R^2	N
1986Q03-2003Q04	-0.043*** (0.130)	1.113 (0.316)	***	-	0.142	70
1999Q02-2003Q04	-0.031 (0.083)	-1.082 (2.562)	-	-	-0.043	19
1990Q02-2003Q04	-0.047 (0.029)	0.151 (0.709)	-	-	0.018	55

Panel B – Annual frequency

	α	β	Reject H_0: β=0	Reject H_0: β=1	Adj R^2	N
1986-2003	-0.053 (0.040)	2.734 (1.922)	-	-	0.105	18
1990-2003	0.048 (0.044)	-0.753 (1.954)	-	-	-0.075	14

Notes:

1. Newey-West standard errors in parentheses). * (**) [***] indicates significance at the 10 (5) [1] per cent level.

   Equation estimated:

   $$s_{t+1} - s_t = \alpha + \beta(i_t^{NZ} - i_t^{AU}) + \epsilon_{t+1}$$

   k =20 for quarterly data

   k =5 for annual data

2. The yield over the five-year horizon is computed as $i = \log(1 + 5 * x)$ where x represents the annual yield in decimal form.

3. The last daily observation of the year is taken to be representative for the whole year for annual data.

Table 7
10-year government bonds

Panel A – Quarterly frequency

	α	β	Reject H_0: β=0	Reject H_0: β=1	Adj R^2	N
1986Q03-1998Q04	-0.053*** (0.130)	0.371 (0.232)	-	***	0.067	50
1990Q02-1998Q04	-0.035** (0.015)	0.759 (0.206)	***	-	0.211	35

Panel B – Annual frequency

	α	β	Reject H_0: β=0	Reject H_0: β=1	Adj R^2	N
1986-1998	-0.071 (0.040)***	0.542 (0.330)	-	-	0.092	13
1990-1998	-0.049 (0.017)**	0.960 (0.270)	***	-	0.390	9

Notes:

4. Newey-West standard errors in parentheses). * (**) [***] indicates significance at the 10 (5) [1] per cent level.

   Equation estimated:

$$s_{t+1} - s_t = \alpha + \beta(i_t^{NZ} - i_t^{AU}) + \epsilon_{t+1}$$

   k =40 for quarterly data

   k =10 for annual data

5. The yield over the ten-year horizon is computed as $i = \log(1 + 10 * x)$ where x represents the annual yield in decimal form.

6. The last daily observation of the year is taken to be representative for the whole year for annual data.

Table 8
Assessing the impact of monetary policy on the exchange rate-interest rate linkage

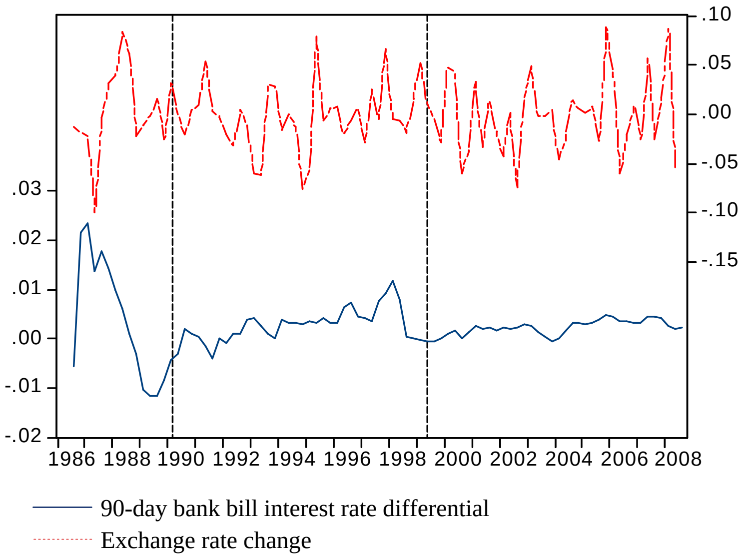
λ_1	$-\frac{1}{\lambda_1\gamma}+\frac{1}{\gamma}$ *γ=0.4*	$-\frac{1}{\lambda_1\gamma}+\frac{1}{\gamma}$ *γ=0.15*
1	0	0
2	1.25	3.33
4	1.875	5
10	2.25	6
50	2.45	6.53

Note:

λ_1 = policy parameter which measures central bank's response to CPI inflation rate deviations from target.

γ = degree of openness (weight on price of imported consumption good in the CPI).

Figure 3
The 90-Day Bank Bill Rate and Exchange Rate Changes



Note: The exchange rate change is expressed as the difference of the log of the quarterly bilateral exchange rate in period *t+1* and the log of same in period *t* while the quarterly interest rate differential is expressed as the log of $(1+\frac{x_t^{NZ}}{4})$ minus the log of $(1+\frac{x_t^{AU}}{4})$. The country-specific annual decimal yield is denoted x_t.