Nominal Exchange Rates, Commodity Prices
and Central Bank Policy

by

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B.Ec.(Hons), Australian National University (1994)

Submitted to the Department of Economics
in partial fulfillment of the requirements for the degree of

Doctor of Philosophy in Economics

at the

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September 2002

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Abstract

This thesis consists of three independent chapters on nominal exchange rates.

The first chapter adds to the forward bias puzzle by noting that while the exchange rate of a small commodity-exporting economy can be closely tied to commodity prices, a portfolio of commodity futures exhibits little if any bias. This is demonstrated for Australia. Using a dependent economy model in which the exchange rate is a function of export prices, three potential explanations for the bias of exchange rate futures, but not commodity futures, are considered. Peso problems do not seem capable of explaining the puzzle. Monetary policy could explain some of the bias, though unlikely the full extent. Systematic expectation errors about the monetary process, while requiring strong assumptions, receive some empirical support from the behaviour of the exchange rate.

The second chapter attempts to resolve the endogeneity of exchange rates and central bank intervention. Using a change in Reserve Bank of Australia intervention policy for identification, simulated GMM is used to estimate a model that includes the contemporaneous impact of intervention. Intervention is found to have an economically significant contemporaneous effect. A $US100m purchase of the domestic currency will appreciate the exchange rate by 1.35 to 1.81 per cent. Further, intervention is found to have the majority of its impact during the day in which it is conducted, with a smaller effect on subsequent days. Australian central bank intervention policy is confirmed to be characterised by leaning against the wind.

The third chapter estimates the dependent economy model outlined in Chapter 1 for the Australian, Canadian and New Zealand dollars. The model provides a good representation of the exchange rates for all three countries up to 1995. In out-of-sample projections the Australian and Canadian models outperform a random walk. The New Zealand model breaks down during the Asian crisis. Commodity futures are used to construct forecasts of the Australian dollar, which at horizons of around one year are more accurate than no-change forecasts.

Thesis Supervisor: Roberto Rigobon
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Thesis Supervisor: Jaume Ventura
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for Nadia
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Grad school is a lot of hard work and a lot of fun. There are many people I'd like to thank who helped make it the rewarding experience it has been.

This thesis is dedicated to the memory of Rudi Dornbusch who has had a significant influence on my work. His teaching stimulated my interest in the topics covered in this thesis. As my advisor, when I was starting out on the treacherous path of research, Rudi directed me away from deadend or shortsighted topics and strongly encouraged me to pursue the topics I ended up researching. If they are of interest, and they have been for me, then it is in part due to Rudi's great insight. Sadly Rudi's declining health prevented him from signing off on my thesis, but in some way he guided it through to completion.

Roberto Rigobon has been an inspirational advisor and co-author. Every time I'd go to ask questions we'd end up taking a walk to Au Bon Pain. But not even that amount of caffeine could explain his unbounded energy and enthusiasm. I'm deeply indebted to Roberto for his support and encouragement.

The comments of Jaume Ventura helped me make the framework I've used more rigorous. He also kindly took the place of Rudi in signing off on my thesis. Guido Kuersteiner's seemingly limitless knowledge of time series econometrics was of great use in deciding on the right estimation strategy for my third paper. Thanks also to Gary King for making life as a grad student at MIT so much easier. I'd also like to thank the Reserve Bank of Australia for funding.

Many of my classmates have contributed to my work indirectly through discussions and keeping me sane. Most of all they've made my time in Cambridge a blast and opened my world to a host of new ideas and cultures. In particular, I'd like to thank Tobias Adrian, Manuel Amador, Kevin Cowan, Toan Do, Francesco Franzoni, Ty Harris, Andrew Hertzberg, John Romalis, Marko Terviö and James Vickery. I'd also like to thank James for a fun living environment and his culinary skills at the Hampshire St palace. All the people in the ANZ club, those who've played touch football and soccer, and the bands who played at the Kendall Cafe and elsewhere have also contributed to Cambridge being a great four years. Many people outside of Cambridge have also helped to make my time in the USA a great experience. My brother Philip, his wife Sid and their gorgeous girls Clare and Isabella in New York have been exceptionally generous and have helped to keep home sickness at bay. Thanks to Nick de Roos,
Bruce Preston and Tasha Rogers for many fun weekends away. And to all my family and friends back home, thanks for your support and encouragement. I can't wait to have a VB sitting on the beach (yes, alcohol outside).

One person has contributed more than any other to me finishing my Ph.D. with a smile on my face. Nadia, I thank you for your patience, perseverance, support, and encouragement. We've proved 10,000 miles is nothing when you're close enough. When we're sitting on our balcony in Bronte looking back over our lives, the months apart will pale next to the years together.
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Introduction

*Markets look a lot less efficient from the banks of the Hudson than from the banks of the Charles*

*Fisher Black*

Commodity currencies: why are exchange rate futures biased if commodity futures are not?

The first chapter adds to the puzzle of the forward bias of exchange rates by noting that while the exchange rate of a small commodity-exporting economy can be closely tied to commodity prices, a portfolio of commodity futures exhibits little if any bias. A dependent economy model is developed in which the exchange rate depends on export prices, as well as import prices, non-traded output and the domestic money supply. The close relationship between commodity prices and the exchange rate is demonstrated for Australia. The bias of exchange rate forwards is shown by the negative coefficient from a 'Fama regression'. This paper finds that the coefficient from an equivalent regression using a portfolio of commodity futures designed to replicate Australian export prices, and so the exchange rate, is positive. The exchange rate model is used to examine whether the domestic monetary stock could cause the bias in exchange rate forwards when there is an absence of bias in commodity futures. Three potential explanations are considered. Peso problems do not seem capable of explaining the puzzle. Monetary policy could explain some of the bias, though unlikely the full extent. Systematic expectation errors about the monetary process, while requiring strong assumptions, receive some empirical support from the behaviour of the exchange rate.
Identifying the efficacy of central bank intervention

The second chapter, co-authored with Roberto Rigobon, considers the endogeneity of exchange rates and intervention that has long plagued studies of the effectiveness of central banks' actions in foreign exchange markets. Researchers have either excluded contemporaneous intervention, so that their explainers are predetermined, or obtained a small, and typically incorrectly signed, coefficient on contemporaneous intervention. Failing to account for the endogeneity, when central banks lean against the wind and trade strategically, will likely result in a large downward bias to the coefficient on contemporaneous intervention – explaining the negative coefficient frequently obtained. Using an alternative identification assumption, a change in Reserve Bank of Australia intervention policy, this paper estimates, using simulated GMM, a model that includes the contemporaneous impact of intervention. There are three main results. The point estimates suggest that central bank intervention has an economically significant contemporaneous effect. A $US100m purchase of the domestic currency will appreciate the exchange rate by 1.35 to 1.81 per cent. This estimate is remarkably similar to the calibration conducted by Dominguez and Frankel (1993c), who themselves noted their estimate was larger than previous empirical findings. Secondly, the vast majority of the effect of an intervention on the exchange rate is found to occur during the day in which it is conducted, with only a smaller impact on subsequent days. Finally, Australian central bank intervention policy is confirmed to be characterised by leaning against the wind.

Modelling commodity currencies: 'The Aussie and two birds'

The third chapter estimates a dependent economy model of the exchange rate for the Australian, Canadian and New Zealand dollars. In this model the exchange rate is determined by balance of payments equilibrium so that the fundamentals of the exchange rate are trade prices, as well as the domestic money supply and non-traded output. The model provides a good representation of the exchange rates for all three countries up to 1995. For Australia and New Zealand the commodity price, serving as a proxy for export prices, is a significant determinant of the exchange rate. However, the Canadian dollar is not a pure 'commodity currency' as a broader export price index is required, rather than the commodity price index, for a well specified model.
In out-of-sample projections after 1995 both the Australian and Canadian models outperform a random walk. The New Zealand model breaks down around the time of the Asian crisis demonstrating some fragility of the model to financial shocks. An index of commodity futures prices is used to construct pure forecasts of the Australian dollar using only past and current information. At horizons of around one year this model is more accurate than no-change forecasts, showing there is some predictability in exchange rates.
Chapter 1

Commodity currencies: Why are exchange rate futures biased if commodity futures are not?

1.1 Introduction

The forward bias of exchange rates has puzzled economists for over two decades. This paper adds to the puzzle by noting that while the exchange rate of a small commodity-exporting economy can be closely tied to commodity prices, a portfolio of commodity futures exhibits little if any bias. The close relationship between commodity prices and the exchange rate is demonstrated for Australia, a country for which approximately 60 per cent of exports are commodities.¹ The bias of exchange rate forwards has typically been demonstrated by observing that the slope coefficient from a regression of the change in the spot price on the forward premium, the 'Fama regression', is less than unity and often, as in the case of Australia, is negative. This paper demonstrates that the coefficient from an equivalent regression using a portfolio of commodity futures is.

¹ Australia is used as futures and forward contracts exist for a large proportion of its exports. Chapter 3 models three prominent 'commodity currencies', the Australian, Canadian and New Zealand dollars and finds a strong relationship between commodity prices and the Australian dollar. A smaller proportion of Canadian exports are commodities and as a result a broader export price index is needed to explain the exchange rate. A large proportion of New Zealand exports are commodities but the relationship between the exchange rate and commodity prices is less stable and fewer of New Zealand's, largely dairy and lamb, commodity exports are covered by futures.
futures designed to replicate Australian export prices, and so the exchange rate, is positive and significant.\footnote{The forward premium regression conducted using an index constructed to replicate Canadian commodity exports also had a positive and significant beta coefficient indicating this result is not peculiar to Australia. However, the Canadian series is not used in this study as the Canadian exchange rate has a weaker relationship with commodity prices as shown in Chapter 3.} This implies that the bias in exchange rate futures must come from one of the other determinants of the exchange rate. This paper shows that in addition to the commodity export price, the exchange rate is also determined by the domestic money stock. The practice of monetary policy or expectation errors about the money stock could cause the bias of exchange rate forwards despite the absence of bias in commodity futures.

The remainder of this chapter proceeds as follows. Section 1.2 outlines the puzzle in more detail and discusses possible explanations. A dependent economy model of a 'commodity currency' is then developed from micro foundations in Section 1.3. This model is used, in Section 1.4, to consider three explanations of the forward bias of exchange rates that depend on the money stock and so could be consistent with unbiased commodity futures. Section 1.4.1 shows how systematically biased expectations regarding the monetary process can cause the puzzle. Alternative explanations deriving from peso problems and the practice of monetary policy are presented in Sections 1.4.2 and 1.4.3. This is followed by some concluding remarks.

1.2 The puzzle

Nominal exchange rates are notoriously difficult to model as demonstrated by Meese and Rogoff (1983) and numerous researchers in their wake.\footnote{See for example the forthcoming February 2003 issue of the Journal of International Economics which publishes papers from the September 2001 conference at the University of Wisconsin marking 20 years since the Meese and Rogoff paper.} Yet the price of commodities is often used as a guide to the value of the exchange rate of commodity exporting countries, at least within financial markets. Figure 1-1 demonstrates this relationship for Australia.\footnote{This graph is drawn using an index of spot prices for commodity futures that is described below.} This is obviously an incomplete representation of the fundamental value of a currency. But it is a crucial component. In the simplest case, consider a small country that exports one good in exchange for a consumption good, and has no access to international borrowing. An increase in the price of the exported good improves the country's terms of trade and so increases their consumption of
the imported good. If monetary policy does not respond to this shock the domestic price of the imported consumption good will fall, which under the Law of One Price (LOP) will imply a depreciation of the exchange rate. This implies an increase in the price of the exported good is associated with a depreciation of the exchange rate, as Figure 1-1 shows is the case for Australia. In this example the money supply is a crucial determinant of the domestic price of the imported good and so the exchange rate. It could feasibly alter the relationship between the exchange rate and commodity price. As argued in this paper, the behaviour of money appears to be the most likely explanation of the differing degrees of bias in exchange rate and commodity futures. Section 1.3 formally develops from micro foundations this model of a 'commodity currency' in which the economy is a price taker in international markets.

So long as covered interest parity holds, which it does for major currencies, the forward bias of exchange rates will be equivalent to the failure of UIP. Since UIP is a central tenet of both empirical and theoretical international macro models the forward bias then has significant implications well beyond the efficiency of the exchange rate market. The bias of forward exchange
rates has been demonstrated in many studies using a simple regression from Fama (1984). The conditional efficiency of exchange rate futures is tested by the regression

\[ s_{t+1} - s_t = \alpha + \beta (f_t - s_t) + \varepsilon_t \]  

(1.1)

where \( s_t \) is the (log) spot exchange rate, and \( f_t \) is the (log) one period forward exchange rate. Conditional efficiency of exchange rate forwards is equivalent to the hypothesis that \( \beta = 1 \).\(^6\) Almost without exception across country pairs and sample periods, this null hypothesis has been rejected, and typically not by a small margin. In many cases the point estimate of \( \beta \) is negative. The Australian dollar is no different. Over the period during which the Australian dollar has floated, the \( \beta \) from this regression using 3-month forwards is -0.94, as shown in Table 1.1. It is not only significantly different from unity but, at the 10 per cent level, significantly less than zero. Conversely, the \( \beta \) from an equivalent 'Fama regression' using indices of the spot and futures prices of commodities, which are described below, is positive, 0.55, and significantly different from zero at the five per cent significance level. Given the economic relationship these series should have, and is borne out in Figure 1-1, this result implies that the forward bias of exchange rates must be attributable to the other determinants of the exchange rate. Three possible explanations are discussed below, but as already suggested, domestic money will be crucial to these explanations. In common with other studies, the point-estimate of \( \beta \) from the exchange rate forward premium regression is found to vary substantially within the sample, as seen in Figure 1-2. The variability in the exchange-rate beta contrasts with the stability of the beta from the commodity regression, which is almost always between zero and two.

1.2.1 The commodity indices

The construction of the commodity indices is outlined briefly here, with greater detail provided in Appendix B.\(^7\) The indices are constructed from portfolios of commodity futures and forwards

\(^5\)The literature on the exchange rate forward bias is truly voluminous and so only incomplete attention is devoted to it here. For surveys see Hodrick (1987), Lewis (1995) and Engel (1996).

\(^6\)Strictly speaking conditional efficiency also implies \( \alpha = 0 \) though this is typically omitted from testing.

\(^7\)There are two traded futures on commodity indices, the CRB futures index and the Goldman Sachs commodity index, but both of these futures have too short a time series for use in this study. Further, their composition does not reflect the export base of Australia and Cashin et al (1999) suggest that index composition does matter for commodity price indices due to commodity specific shocks.
Table 1.1: Forward Premium Regressions: Exchange rates and commodity prices.

<table>
<thead>
<tr>
<th>Dependent variable: 3-month change in spot price of:</th>
<th>exchange rate</th>
<th>commodity price index</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>full index</td>
<td>futures only</td>
</tr>
<tr>
<td>constant</td>
<td>0.02</td>
<td>0.00</td>
</tr>
<tr>
<td></td>
<td>(0.01)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>forward premium</td>
<td>-0.94</td>
<td>0.55</td>
</tr>
<tr>
<td></td>
<td>(0.58)</td>
<td>(0.21)</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.018</td>
<td>0.046</td>
</tr>
<tr>
<td>Durbin Watson</td>
<td>0.66</td>
<td>0.60</td>
</tr>
</tbody>
</table>

Exchange rate sample is 1/1984 - 9/2000
Commodity price samples are 11/1983 - 9/2000

Figure 1-2: Betas from rolling 5-year forward premium (Fama) regressions.
trading in US dollars on the London Metal Exchange (LME) and various commodity exchanges in the USA. The index weights correspond to the export-share weights from the Reserve Bank of Australia (RBA) commodity price index. Two indices are constructed, an index of commodity spot prices and an index of 3-month futures prices. A 3-month horizon is chosen as the LME forwards are available for 3-month contracts, as are most futures. There are seventeen traded contracts in the indices, covering just over half of the RBA index. As is typically done when working with commodity futures, spot prices are constructed from expiring contacts to avoid problems of basis risk when comparing the futures and spot indices. Because not all futures contracts expire each month, prices in missing months are extrapolated from longer horizon contracts, following Pindyck (1993). This clearly introduces some noise, but the behaviour of the prices of forwards, for which 3-month contracts are available each month and so no extrapolation is needed, serves as a robustness check. As shown in Table 1.1, regressions using sub-indices which separate the futures and forwards have almost identical coefficients on the forward premium, 0.56 and 0.75 respectively, and again both are significantly greater than zero. The large and significant coefficient when using just commodity forwards indicates that while the extrapolation used to construct missing months for commodity futures prices likely contributes to attenuation bias, it is not driving the finding that the bias in the commodity futures index is at most small. Notably, the coefficient from the regression using only futures is insignificantly different from unity so that the hypothesis that this sub-index is conditionally unbiased cannot be rejected.

Despite its smaller coverage, the spot futures commodity price index does an excellent job of replicating the movements in the comprehensive RBA index. The spot commodity price index has a 0.95 correlation with the RBA index, and correlation of 0.59 with the exchange rate, as can been seen in Figure 1-1. The properties of the forward premiums and 3-month spot price returns for the exchange rate and commodity price index are shown in Table 1.2. The commodity price index is seen to have similar volatility to the exchange rate, but notably the commodity index forward premium, \( f_t - s_t \), is more volatile than the exchange rate forward

---

8 There are more futures than forwards in the index and so it is referred to as a 'commodity futures index'.

9 This is only slightly lower than the 0.65 correlation between the RBA index and the exchange rate. The similar movements in the futures index and the broader RBA index are likely in part contributed to by the common determinants of commodity prices (Borensztein and Reinhart (1994)). Pindyck and Rotemberg (1990) suggest that commodity prices display excess comovement, though Cashin et al. (1999) dispute this.
premium. Similarly it is less autocorrelated. These summary statistics are not dissimilar when considering futures and forwards separately.

Table 1.2: Properties of Forward Premiums and Spot Returns for both the Exchange Rate and Commodity Price.

<table>
<thead>
<tr>
<th></th>
<th>exchange rate</th>
<th>commodity price</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>change</td>
<td>forward</td>
</tr>
<tr>
<td>mean</td>
<td>0.81</td>
<td>0.79</td>
</tr>
<tr>
<td>standard deviation</td>
<td>5.37</td>
<td>0.87</td>
</tr>
<tr>
<td>standard deviation / mean</td>
<td>6.65</td>
<td>1.10</td>
</tr>
<tr>
<td>autocorrelation</td>
<td>-0.09</td>
<td>0.87</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th></th>
<th>commodity price</th>
<th>commodity price</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>futures only</td>
<td>forwards only</td>
</tr>
<tr>
<td>change</td>
<td>change</td>
<td>forward</td>
</tr>
<tr>
<td>mean</td>
<td>-0.17</td>
<td>0.34</td>
</tr>
<tr>
<td>standard deviation</td>
<td>4.59</td>
<td>2.16</td>
</tr>
<tr>
<td>standard deviation / mean</td>
<td>-27.01</td>
<td>6.44</td>
</tr>
<tr>
<td>autocorrelation</td>
<td>0.02</td>
<td>0.40</td>
</tr>
</tbody>
</table>

1.2.2 Explaining the bias

The rejection of conditional unbiasedness for exchange-rate forwards, \( \beta \neq 1 \), implies that there are predictable excess returns. This is demonstrated by rewriting equation (1.1) as

\[
s_{t+1} - f_t = \alpha + (\beta - 1) (f_t - s_t) + \varepsilon_t
\]

(1.2)

With \( \beta \neq 1 \) the forward premium, \( f_t - s_t \), which is known at time \( t \), can be used to predict the difference between the realised exchange rate at time \( t + 1 \) and the time \( t \) price at which time \( t + 1 \) currency can be bought and sold. In other words, the forward premium can be used to predict the excess return, \( er_{t+1} = s_{t+1} - f_t \). The identity in equation (1.3) decomposes the predicted excess returns, \( per \), as the sum of a time-varying risk premium, and a market expectation error.
\[ \text{per}_{t+1} = E_t s_{t+1} - f_t = (E_t^{m} s_{t+1} - f_t) + (E_t s_{t+1} - E_t^{m} s_{t+1}) \] (1.3)

where $E_t^{m}$ is the market expectation formed at time $t$, and $E_t$ is the rational expectations operator. The first term in parentheses is the risk premium, while the second is the market expectations error. As many authors have shown drawing on Fama, in order for $\beta < 1$ the predicted excess return must be time varying, and negatively correlated with the forward premium. Even more puzzling, a finding of $\beta < 0$ implies that the predicted excess returns must be more volatile than exchange rate returns. Standard models of risk premia are unable to deliver these properties, as summarised by Engel (1996). Efforts to generate these properties with more complex models, such as first-order risk aversion (Bekeart et. al. (1997)) or habit persistence (Backus et. al. (1993)) have met with a similar lack of success.

The alternative avenue of explanations, market expectation errors, appears to hold more promise in finding an answer to the puzzle. These explanations have the common element that the ex post expectations formed by the researcher differ from those formed ex ante by agents, based on their subjective probability distribution. One possibility is that there are 'peso problems', that some event given non-zero probability weight by agents occurred with an unrepresentative frequency in the researcher's sample. Uncertainty as to whether the event occurred, with Bayesian updating of expectations, can also contribute to seemingly systematic expectation errors that may be correlated with the forward premium. Expectations could also be systematically biased possibly resulting in the exchange rate adjusting slowly to changes in the forward premium, as suggested by Froot and Thaler (1990). Gourinchas and Tornell (2001) provide a tractable specification for biased expectations which demonstrates that the slow adaptation of expectations to monetary factors can result in the forward bias of exchange rates. These explanations could potentially explain the bias of exchange rate forwards, but not of commodity futures, if they are applied to one of the other determinants of the exchange rate, notably the domestic money stock.

A third class of explanations stems directly from the practice of monetary policy, and so could also be consistent with the exchange rate–commodity price bias puzzle. McCallum (1994) demonstrates that in the presence of a persistent deviation from UIP (which need not be a risk premium) monetary policy aimed at minimising exchange rate changes can result in a negative
covariance between the forward premium and predicted excess returns, and so a downward biased beta. An alternative explanation relating to policy is proposed by Fisher Black in Engel (1996). He is quoted as suggesting that central bank intervention could be responsible for the forward bias of exchange rates. But the consistency of the finding of forward bias across countries and time periods, while the practice of intervention has been anything but consistent, suggests this is unlikely to be the cause. Notably, two similar countries Australia and New Zealand have equivalent degrees of exchange rate forward bias despite the Australian central bank frequently intervening in the foreign exchange market while the New Zealand central bank has never done so.

Commodity markets are in many ways more complex than foreign exchange markets, due to the existence of storage for many commodities, seasonal factors and illiquid markets. In particular because of the physical nature of the goods traded in commodity markets there are different influences on risk premia. An extensive literature exists relating futures risk premia to systematic risk and hedging pressures, for example see Bessembinder (1992) and Hirshleifer (1988, p39). But while the existence of inventory means that spot and futures prices will be codetermined, as modeled by Hong (2000) and Routledge et al (2000), the existence of inventory does not necessarily infer that commodity futures need be either biased or unbiased predictors of future spot prices. Just as time varying risk premia, expectation errors or peso problems could be responsible for the bias in exchange rate markets, these factors could be present in commodity markets. Fama and French (1987) consider the determinants of futures prices for 21 individual commodities. They find that for each of the commodities, the point estimate of $\beta$ from a regression of the form of equation (1.1) is between zero and one, though because of large standard errors it is significantly greater than zero in only seven cases. Despite this evidence of time-varying predictable returns, they find a greater role for the theory of storage in the formation of commodity futures prices than for time varying risk premia. By combining

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10This is much broader than the question, dating back to Keynes, as to whether futures prices must rise as they approach expiration, that is they demonstrate normal backwardation, to compensate speculators. After comprehensive testing Kolb (1992) suggests normal backwardation isn't so normal. Normal backwardation would imply $\alpha \neq 0$ but not necessarily that $\beta \neq 0$.

11The 'theory of storage' of Kaldor (1939), Tesler (1958) and Working (1949) describes the futures price of a storable commodity, for which inventory is held, as a function of the interest foregone, warehousing costs and convenience yield (the benefit from holding the physical commodity rather than a futures contract). The impact storage can have on the dynamics of the spot price is estimated by Deaton and Laroque (1996).
commodities into portfolios many of the idiosyncratic factors that affect the cost of storage for individual commodities can be mitigated to a large extent. That is the approach pursued in this paper.

1.3 Dependent economy model of the exchange rate

Dependent economy models, for example Dornbusch (1980), consider small countries that are assumed to be unable to influence conditions in the rest of the world. This framework is especially relevant for small commodity exporting countries, such as Australia, whose export product is homogenous and market share is small.\footnote{The intuition of this framework has inspired many empirical studies. Freebairn (1990) uses this intuition to motivate his study of the relationship between the Australian dollar and commodity prices. Similar analysis was performed for the real exchange rate in Gruen and Wilkinson (1994). Sjaastad (1998) took the dependent economy intuition beyond commodity exporting countries to examine the behaviour of the Swiss franc.}

This section develops a simple dependent economy model of the nominal exchange rate using micro foundations. The economy is endowed with a fixed quantity per period, normalised to unity, of a good, $X$, that is exported and not consumed. The assumption of a fixed supply is appropriate for a commodity exporting country; the large investments for mineral extraction and the delay to harvest of agricultural commodities imply that their supply is inelastic. The world price of the export good, $P^*_{Xt}$, is taken as given and assumed to be a log normal random walk.

$$\log P^*_{Xt+1} \sim N \left( \log P^*_{Xt}, \sigma^2_{P_X} \right)$$

This assumption contradicts the finding in Section 1.1 that there are predictable changes in commodity prices, but is made to focus the attention on the bias present in exchange rate futures. Standard notation of a star to indicate the value taken by a foreign variable, and lower case variables to indicate the log of their uppercase equivalents, will be used throughout. All variances are assumed to be constant through time. The country imports a traded good for consumption, $T$, whose world price is also assumed to follow a log normal random walk. Agents also consume a non-traded good, $Y_{Nt}$, of which the economy receives a random per period endowment. This too follows a log normal random walk. Preferences are a function of
aggregate consumption and real money balances and are given by

\[ U_t = \sum_{s=t}^{\infty} \beta^{s-t} \left[ \frac{C_s^{1-\rho}}{1-\rho} + \frac{\chi}{1-\delta} \left( \frac{M_s}{P_s} \right)^{1-\delta} \right] \]  \hspace{1cm} (1.4)

where \( C_t \) is a Cobb-Douglas aggregation over the traded good and the non-traded good, \( C = C_N^\gamma C_T^{1-\gamma} / \gamma (1-\gamma)^{1-\gamma} \). The representative consumer maximises the expectation of (1.4) subject to the budget constraint

\[ B_{t+1} P_{t+1} + \tilde{B}_{t+1} + S_t \tilde{B}_{t+1}^* + M_t \]

\[ = (1+r) B_t P_{t+1} + (1+i_t) \tilde{B}_t + (1+i_t^*) S_t \tilde{B}_t^* + M_{t-1} + P_{Xt} + Y_{Nt} P_{Nt} - C_t P_t - \tau_t P_t \]

where \( P_t = P_{Nt}^\gamma P_{Tt}^{1-\gamma} \) is the economy price index; \( B_t, \tilde{B}_t, \) and \( \tilde{B}_t^* \) are holdings at the start of period \( t \) of the real bond, denominated in the tradable good and paying a constant rate of return \( r \), the domestic nominal bond, and the foreign nominal bond; \( S \) is the exchange rate (price of foreign currency); and, \( M_{t-1} \) is the holdings of money at the start of period \( t \). The bonds, of which initial holdings are zero, are the only financial assets; there are not complete markets. Since there is a representative consumer and money has no real effects Ricardian equivalence will hold and so the value of government transfers, \( \tau_t P_t \), will cancel with the two nominal money holding terms in the budget constraint. There are no restraints to trade or transport costs so that the LOP holds, \( P_T = S P_T^* \), which will define the exchange rate. Prices are flexible in the model, although the implication of sticky prices is discussed.

The assumption of random walk exogenous prices and the real-nominal dichotomy produces a neat closed form for the model without any assumptions on the behaviour of, or the nature of expectations about, the money supply. This basic model then serves as the building block to consider different potential causes of the forward bias of exchange rate futures.

Since the model structure is standard, details of the solution are relegated to Appendix A, and only a sketch is outlined here. All exogenous variables are log normal and so the solution will have a log normal form. The first order conditions imply that the paths of tradable and non-tradable consumption are governed by

\[ \gamma (\rho - 1) E^m_t \{ c_{Nt+1} - c_{Nt} \} + (\rho (1-\gamma) + \gamma) E^m_t \{ c_{Tt+1} - c_{Tt} \} = K_1 \]  \hspace{1cm} (1.5)
where \( E_t^m \) is the market expectation operator and the letter \( K \) with a subscript is used to indicate constants (that are functions of the exogenous parameters). Consumption of the non-traded good must equal its production, which is a log random walk. This condition then confirms that consumption of the traded good will also be log normally distributed. As a result of the log random walk of exogenous prices, trade will be balanced in each period.\(^{13}\)

The derivation of the exchange rate starts with the first order condition with respect to money balances.

\[
1 - \chi \left( \frac{M_t}{P_t} \right)^{-\delta} C_t = \beta E_t^m \left\{ \left( \frac{C_t}{C_{t+1}} \right)^\rho \frac{P_t}{P_{t+1}} \right\} \quad (1.6)
\]

The left hand side of this expression is not log linear, and so is linearised around the non-stochastic steady state. Using the first order condition with respect to domestic nominal bonds, in the steady state \( 1 - \chi \left( \frac{M}{P} \right)^{-\delta} \bar{C}^\rho = \frac{1}{1+i} \) where an overbar indicates the value of a variable in equilibrium. The log linearisation of the left hand side of equation (1.6) is then

\[
\log \left( 1 - \chi \left( \frac{M_t}{P_t} \right)^{-\delta} C_t \right) = \tilde{i} \delta (m_t - p_t) - \tilde{i} \rho c_t - \tilde{i} \log \left( \frac{\chi}{i} (1 + \tilde{i}) \frac{1+i}{i} \right) \quad (1.7)
\]

Equating this to the log of the right hand side of (1.6) gives

\[
\delta (m_t - p_t) - \rho c_t - \log \left( \frac{\chi}{i} (1 + \tilde{i}) \frac{1+i}{i} \right) = \frac{1}{i} \log \beta - \frac{1}{2i} E_t^m \{p_{t+1} - p_t\} - \frac{\rho}{i} E_t \{c_{t+1} - c_t\} + \frac{1}{2i} \text{var} (p_{t+1} - \rho c_{t+1}) \quad (1.8)
\]

Using conditions already established it is easily shown that \( E_t^m \{p_{t+1} - p_t\} = E_t^m \{s_{t+1} - s_t\} \) and \( E_t^m \{c_{t+1} - c_t\} = 0 \). Substituting these into equation (1.8) delivers

\[
\tilde{i} \delta (m_t - s_t) - \tilde{i} (\delta \gamma + \rho (1 - \gamma)) p_t^* - \tilde{i} (\delta - \rho) (1 - \gamma) p_{t+1}^* + \tilde{i} \gamma (\delta - \rho) y_{Nt} \quad (1.9)
\]

\[
= -E_t^m \{s_{t+1} - s_t\} - K_2
\]

where \( K_2 \) is a constant.\(^{14}\) Substituting forward, and imposing that there are no bubbles, the

\(^{13}\) This assumes constraints on the parameter values that ensure a solution exists as described in the Appendix.

\(^{14}\) \( K_2 = \log \beta + \frac{1}{2} \text{var} (p + \rho c) + \tilde{i} \delta \gamma \log \left( \frac{\chi}{1+i} \right) - \tilde{i} \rho \log (\gamma (1 - \gamma)^{1-\gamma}) + \tilde{i} \left( \frac{\chi}{i} (1 + \tilde{i}) \frac{1+i}{i} \right) \)
exchange rate is derived as

\[
s_t = E^m_t \left\{ \sum_{j=0}^{\infty} \left( \frac{1}{1 + i\delta} \right)^j m_{t+j} \right\} - \left[ \gamma (p^*_{Xt} + y_{Nt}) + \left( 1 - \gamma \right) p^*_{Tt} \right] - \frac{\rho(1 - \gamma)}{\delta} (p^*_{Xt} - p^*_{Tt}) - \frac{\gamma \rho}{\delta} y_{Nt} + K_2 \tag{1.10}\]

This equation highlights two channels through which the world price of the exported good can affect the nominal exchange rate. The real wealth effect is represented by the third term. An increase in the price of the exported good increases the net present value of expected wealth and so the consumption of the imported good. In order to obtain internal balance the domestic relative price of the imported good must fall, so that by the LOP the domestic currency appreciates. This effect is greater the more open is the economy, \( 1 - \gamma \), and the larger the intertemporal elasticity of substitution of consumption, \( \rho \), relative to the elasticity of money demand, \( \delta \).

The second channel occurs through the impact on the general price level and so the real money supply. The second term in equation (1.10) is the domestic price level, measured in units of the foreign currency. An increase in the price of the exported good increases the domestic price level in foreign currency terms. To restore real money demand equilibrium the exchange rate must appreciate.

The exchange rate depends on the future stream of the domestic money supply, which could potentially respond to the exogenous variables. The exchange rate also appreciates in the output of the non-traded sector both through the relative price change from the substitution effect, and the impact on the aggregate price level.

The possibility for sectoral switch of labour is obscured in this model by the exogenous supplies of the exported and non-traded goods. If output were to respond to relative price changes – through labour switching between the sectors – then the effect on the exchange rate of the world price of the exported good would be reinforced. An increase in its price would lead to an increase in the production of the exported good, and a contraction in output of the non-traded good. The relative price of the traded good would fall even further in the home country resulting in a larger appreciation of the exchange rate.
Sticky prices could weaken the commodity price-exchange rate link by reducing the margin on which prices can adjust to attain internal balance. Similarly, more extensive financial markets which lessen the exposure of domestic agents to the world commodity price could weaken the commodity currency relationship. In practice, neither of these are likely to completely break the correlation of the exchange rate with commodity prices.

1.4 Explanations of the puzzle

Section 1.3 showed that in addition to the commodity price, the exchange rate is also determined by the domestic money supply, the non-traded output and the price of imported goods. While there is no futures market for the price of imported goods, given commodity futures prices display little bias it would seem unlikely that the price of the imported good causes the bias in exchange rate forwards. Similarly, there is no evidence of significant expectation errors or peso-type problems with non-traded output. Rather, the behaviour of money seems to be the more likely cause for the bias in exchange rate forwards despite the absence of such bias in commodity futures. Given this, the dependent economy model was developed to provide the maximum flexibility to the behaviour of, and expectations about, the money supply. The model of the exchange rate assumed non-monetary exchange rate determinants are random walks, simplifying the model to focus attention on monetary factors. This obscures the fact that a component of the exchange rate fundamental value is predictable and such predictions are virtually unbiased. In reality the factor inducing negative covariance between the interest differential and realised exchange rate return must be even larger to account for the positive covariance that commodity price changes would induce.

This section uses the dependent economy model to consider three leading explanations of the forward bias of exchange rates that rely on the money stock and so could be consistent with the lack of bias in commodity futures. Section 1.4.1 uses the model of Gourinchas and Tornell (2001) to demonstrate how systematically biased expectations about monetary policy could cause the observed relationships. The possibility that peso problems are to blame is considered in Section 1.4.2 while Section 1.4.3 considers whether the objective of monetary policy could drive the result.
1.4.1 Systematic expectation errors

Gourinchas and Tornell (2001) attribute the observed forward bias to agents' perception that interest rate shocks contain a large transitory component when in reality they are highly persistent. The initial under estimation, and gradual learning, of the change in the future fundamentals of the exchange rate results in a gradual response to an interest rate shock. The gradual appreciation (depreciation) of the exchange rate as agents learn about the persistence of a positive (negative) interest rate shock results in a downward bias in the 'Fama coefficient'. Gourinchas and Tornell justify their expectations theory by demonstrating that expectations of interest rates from surveys do not reflect the true persistence in interest rates.

Here Gourinchas and Tornell's expectations misperception is adapted to apply directly to money in order to fit in with the model of the exchange rate developed in Section 1.3. Since Section 1.1 demonstrated that there is very little bias in commodity futures, in order to focus on explanations of the bias in exchange rate forwards, it is assumed that agents have rational expectations with regard to foreign prices. This is a strong assumption but is shown to be consistent with the behaviour of the exchange rate. Implicit in this assumption is that foreign prices are determined in such a way that any misperception about foreign money shocks does not carry over to foreign prices.

So long as agents understand the relationship between fundamentals and the money supply, changes in the money supply due to fundamentals can not introduce the bias in forward exchange rates. To simplify notation then, the money supply is assumed to be unrelated to economic fundamentals. The money supply is proposed to follow a persistent process

\[
m_t = \lambda m_{t-1} + \epsilon_t \tag{1.11}
\]

\[
= \sum_{j=0}^{\infty} \lambda^j \epsilon_{t-j} \tag{1.12}
\]

where the constant term is normalised to zero. Agents believe that there is a larger transitory component to shocks than in reality. Specifically they believe that the money process is

\[
m_t = z_t + v_t \tag{1.13}
\]

\[
z_t = \lambda z_{t-1} + \epsilon_t
\]

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where their misperception is governed by the value of $\sigma_v^2 = \text{var}(v)$. If $\sigma_v^2 = 0$ then expectations reflect the true money process. As described in Gourinchas and Tornell, agents will use a Kalman filter to produce forecasts of money so that their expectations are given by

$$E_t^m m_{t+1} = k \lambda m_t + (1-k) \lambda E_{t-1}^m m_t$$  \hspace{1cm} (1.14)$$

where $k \in [0, 1]$ is a constant that depends on the relative sizes of the variances $\sigma_v^2$ and $\sigma_e^2$ and other model parameters. The constant $k$ governs the degree of bias in expectations, the larger is $k$ the less is the bias, with rational expectations corresponding to $k = 1$. Expression (1.14) is derived in Appendix A. From equation (1.14) it follows that $E_t^m m_{t+j} = \lambda^{j-1} E_t^m m_{t+1}$. This can then be used to simplify the exchange rate given by equation (1.10) to

$$s_t = \frac{i^d}{1 + i^d} \left( m_t + \frac{1}{1 + i^d - \lambda} E_t^m m_{t+1} \right) - \left[ \gamma p^*_{X_t} + (1 - \gamma) p^*_{T_t} \right] - \frac{\rho (1 - \gamma)}{\delta} (p^*_{X_t} - p^*_{T_t}) - \frac{\gamma (\delta + \rho)}{\delta} y_{N_t} + K_2$$  \hspace{1cm} (1.15)$$

As shown in Section 1.1 the bias in $\beta$, the 'Fama coefficient', can be calculated from the covariance of the market expectation error with the interest differential. As a first step to calculating these expressions it is useful to first express the expected money stock in terms of past shocks. The expected value of future money will be a discounted sum of past values of the money supply, and so a discounted sum of past shocks. By recursively substituting equation (1.14) and then substituting in equation (1.12) the expected money supply is given by

$$E_t^m m_{t+1} = k \lambda \sum_{j=0}^{\infty} \left\{ \left[ k (1 - \lambda) \right]^j \sum_{i=0}^{\infty} \lambda^i \varepsilon_{t-j-i} \right\}$$

$$= \sum_{j=0}^{\infty} \lambda^{j+1} \left[ 1 - (1-k)^{j+1} \right] \varepsilon_{t-j}$$  \hspace{1cm} (1.16)$$

The misperception as to the persistence of shocks, $k < 1$, leads agents to underestimate the impact of shocks on the future money supply. Variables other than money are expected to follow random walks so the expected depreciation depends only on the current, and next period’s expected, money supplies. Both of these variables are a weighted sum of the history.
of shocks to money. The depreciation expected by the market can then be calculated as

\[
E_t^m \{s_{t+1} - s_t\} = \frac{\bar{i}\delta}{1 + i\delta} \left[ \left( E_t^m m_{t+1} + \frac{\lambda E_t^m m_{t+1}}{1 + i\delta - \lambda} \right) - \left( m_t + \frac{1}{1 + i\delta - \lambda} E_t^m m_{t+1} \right) \right]
\]

\[
= \frac{\bar{i}\delta}{1 + i\delta} \left[ \frac{\bar{i}\delta}{1 + i\delta - \lambda} \sum_{j=0}^{\infty} \lambda^j \left[ 1 - (1 - k)^{j+1} \right] \varepsilon_{t-j} - \sum_{j=0}^{\infty} \lambda^j \varepsilon_{t-j} \right]
\]

\[
= \frac{-\bar{i}\delta (1 - \lambda)}{1 + i\delta - \lambda} \sum_{j=0}^{\infty} \lambda^j \varepsilon_{t-j}
\]

\[
- \frac{(i\delta)^2 \lambda (1 - k)}{(1 + i\delta)(1 + i\delta - \lambda)} \sum_{j=0}^{\infty} \lambda (1 - k)^j \varepsilon_{t-j}
\]

(1.17)

The first term in equation (1.17) is the expected depreciation in a world with rational expectations and results from the persistence of monetary shocks. The second term introduces the impact of systematic expectation errors. The expression for the expected depreciation can then be substituted into the UIP expression derived from the first order conditions with respect to holdings of domestic and foreign nominal bonds.

\[
i_t - i_t^* = E_t^m \{s_{t+1} - s_t\} + \frac{1}{2} \text{var} (s) + \text{cov} (s, \rho c + p)
\]

(1.18)

\[
= \frac{-\bar{i}\delta (1 - \lambda)}{1 + i\delta - \lambda} \sum_{j=0}^{\infty} \lambda^j \varepsilon_{t-j} - \frac{(i\delta)^2 \lambda (1 - k)}{(1 + i\delta)(1 + i\delta - \lambda)} \sum_{j=0}^{\infty} \lambda (1 - k)^j \varepsilon_{t-j}
\]

\[
+ \frac{1}{2} \text{var} (s) + \text{cov} (s, \rho c + p)
\]

(1.19)

When expectations about the money supply aren’t rational, market expectations of next periods’ exchange rate will be systematically biased. The market expectation error is calculated
\[ 
\zeta_t = E_t s_{t+1} - E_t^m s_{t+1} 
= \frac{i\delta}{1 + i\delta} \left[ \lambda m_t + \frac{\lambda^2 km_t + \lambda (1-k) E_t^m m_{t+1}}{1 + i\delta - \lambda} \right] - \frac{1 + i\delta}{1 + i\delta - \lambda} E_t^m m_{t+1} 
= \frac{i\delta (1 + i\delta - \lambda (1-k))}{(1 + i\delta)(1 + i\delta - \lambda)} [\lambda m_t - E_t^m m_{t+1}] 
= \frac{i\delta (1 + i\delta - \lambda (1-k))}{(1 + i\delta)(1 + i\delta - \lambda)} \left[ \sum_{j=0}^{\infty} [\lambda (1-k)]^{j+1} \varepsilon_{t-j} \right] 
\]

(1.20)

The second line follows from the definition of market expectations in equation (1.14). The market expectation error arises because agents place too little weight on more recent shocks as they believe these are in part transitory. The magnitude of the error declines as the degree of misperception, \(1 - k\), declines.

The 'Fama coefficient' is the coefficient from the regression of the realised change in the exchange rate on the interest differential. Using the fact that the rational expectations error is by definition uncorrelated with time \(t\) variables \(\beta\) can be expressed as

\[
\text{plim} \left( \beta \right) = \frac{\text{cov} \left( i_t - i_t^*, \Delta s_{t+1} \right)}{\text{var} \left( i_t - i_t^* \right)} 
= \frac{\text{cov} \left( i_t - i_t^*, E_t^m \Delta s_{t+1} \right)}{\text{var} \left( i_t - i_t^* \right)} + \frac{\text{cov} \left( i_t - i_t^*, E_t s_{t+1} - E_t^m s_{t+1} \right)}{\text{var} \left( i_t - i_t^* \right)} 
\]

By equation (1.18) the first term is unity. Substituting equations (1.19) and (1.20) into the second term, the probability limit of the coefficient \(\beta\) is given by

\[
\text{plim} \left( \beta \right) = 1 - \left[ \frac{\lambda (1-k)(1 + i\delta - \lambda (1-k))(1 - i\delta)(1-\lambda)}{1 - \lambda^2(1-k)} + \frac{i\delta(1 + i\delta - \lambda (1-k))\lambda^2(1-k)^2}{1 - \lambda^2(1-k)^2} \right] 
= \left[ \frac{(1 + i\delta)^2(1-\lambda)^2}{1 - \lambda^2} - \frac{2i\delta \lambda (1-\lambda)(1-k)}{1 - \lambda^2(1-k)} + \frac{(i\delta)^2 \lambda^2(1-k)^2}{1 - \lambda^2(1-k)^2} \right]^{-1} 
\]

Since \(k, \lambda \in [0, 1]\) the term in brackets will always be greater or equal to zero and so \(\text{plim} \left( \beta \right)\) will be less than or equal to unity. The range of values for \(\beta\), as a function of \(\lambda\) and \(k\), is shown in Figure 1-3. This is drawn for a Taylor expansion value of \(i\delta = 0.05\) but the value of
Figure 1-3: Systematic expectation errors: the 'Fama coefficient' from a regression of the exchange rate return on the forward premium.

\[ \beta \]

\[ \lambda \]

\[ k \]

\( \beta \) is insensitive to this choice, except for large values of \( \lambda \).\textsuperscript{15} Beta is biased downward because agents under-estimate the persistence of monetary shocks and so the exchange rate initially under responds to shocks. The revision to expectations as agents learn the persistence of the shock results in a tendency for the exchange rate to continue its initial trajectory. Figure 1-3 confirms the intuition of the model that \( \beta \) will be more biased, that is larger negative values, for highly persistent monetary shocks (large \( \lambda \)) and greater bias of expectations (smaller \( k \)). For \( \lambda \) very close to unity (larger than the 0.99 shown on the graph) the biased expectations can result in a large negative value for \( \beta \) (even double digit values), though as noted before, for large \( \lambda \), \( \beta \) is sensitive to the value of \( \bar{\delta} \).

\textsuperscript{15} This value is selected for the average monthly interest rate is \( \bar{i} = 0.01 \) and the elasticity of money demand \( \delta = 5 \).
Figure 1-4: Number of periods of overshooting and sign of beta for various parameter combinations.

Another result of the biased expectations is that the exchange rate can demonstrate delayed overshooting, of the form found empirically by Eichenbaum and Evans (1995). The impulse response function of the exchange rate to a money shock is

\[ s_{t+j} = \frac{-i\delta (1 - \lambda)}{1 + i\delta - \lambda} \lambda^j \varepsilon_{t-j} - \frac{(i\delta)^2 \lambda (1 - k)}{(1 + i\delta)(1 + i\delta - \lambda)} \left[ \lambda (1 - k) \right]^j \varepsilon_{t-j} \] (1.21)

Again the first term represents the rational expectations response of the exchange rate to monetary shocks, and the second the impact of biased expectations. If the second decays slowly, because \( k \) is small, then it is possible for the response of the exchange rate to a monetary shock to be larger after several periods than it is immediately. Figure 1-4 shows the range of parameters for which there is overshooting of various lengths.
Testing the model

The persistence of monetary policy, $\lambda$, can be estimated using a Kalman Filter. The results of this estimation, in Table 1.3, show that the money supply is highly persistent, $\lambda$ is essentially unity, and the variance of transitory shocks is found to be zero, consistent with the money process in equation (1.12).

<table>
<thead>
<tr>
<th>$\lambda$</th>
<th>$\bar{m}$</th>
<th>$\sigma_\varepsilon$</th>
<th>$\sigma_v$</th>
</tr>
</thead>
<tbody>
<tr>
<td>0.9999</td>
<td>0.7415</td>
<td>0.0067</td>
<td>0.0000</td>
</tr>
<tr>
<td>(0.0002)</td>
<td>(0.4226)</td>
<td>(0.0003)</td>
<td>(0.0005)</td>
</tr>
</tbody>
</table>

Standard errors in brackets.
The estimated model is:

$m_t = \bar{m} + z_t + \nu_t$
$z_t = \lambda z_{t-1} + \varepsilon_t$

The markets' expectation for monetary policy cannot be measured directly for Australia and so it is not possible to estimate $k$, as Gourinchas and Tornell can do using survey based interest rate expectations. An alternative approach is to infer the values of $k$ that are consistent with the exchange rate overshooting implied by the model. A VAR using monthly data from July 1984 to December 2000 is estimated based on the dependent economy model developed in Section 1.3.16 A VAR estimation is used here to facilitate comparison with previous research such as Eichengreen and Evans (1995).17 Import (traded goods) prices are measured by the price of exports of industrialised countries (since monthly import prices are not available for Australia) and the output in the non-traded sector is proxied by the unemployment rate since this is available monthly and a large portion of non-traded output comes from the labour-intensive service sector. The data are described in Appendix B. The VAR ordering reflects the dependent economy assumption that foreign prices are exogenous, but the exchange rate depends on these prices, non-traded output, and the money supply. The money supply is allowed to depend on

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16 The Australian dollar floated in December 1983. The sample is constrained to start in January 1984 by the availability of the commodity price data. The VAR was tested with up to 6 lags resulting in a start date of July 1984.

17 The results for the VAR are consistent with the single equation approach used in Chapter 3 to estimate commodity currency models.
Testing the model

The persistence of monetary policy, $\lambda$, can be estimated using a Kalman Filter. The results of this estimation, in Table 1.3, show that the money supply is highly persistent, $\lambda$ is essentially unity, and the variance of transitory shocks is found to be zero, consistent with the money process in equation (1.12).

<table>
<thead>
<tr>
<th>$\lambda$</th>
<th>$\bar{m}$</th>
<th>$\sigma_e$</th>
<th>$\sigma_v$</th>
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<td>0.9999</td>
<td>0.7415</td>
<td>0.0067</td>
<td>0.0000</td>
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<tr>
<td>(0.0002)</td>
<td>(0.4226)</td>
<td>(0.0003)</td>
<td>(0.0005)</td>
</tr>
</tbody>
</table>

Standard errors in brackets.
The estimated model is:

$m_t = \bar{m} + z_t + v_t$

$z_t = \lambda z_{t-1} + \varepsilon_t$

The markets' expectation for monetary policy cannot be measured directly for Australia and so it is not possible to estimate $k$, as Gourinchas and Tornell can do using survey based interest rate expectations. An alternative approach is to infer the values of $k$ that are consistent with the exchange rate overshooting implied by the model. A VAR using monthly data from July 1984 to December 2000 is estimated based on the dependent economy model developed in Section 1.3.\textsuperscript{16} A VAR estimation is used here to facilitate comparison with previous research such as Eichengreen and Evans (1995).\textsuperscript{17} Import (traded goods) prices are measured by the price of exports of industrialised countries (since monthly import prices are not available for Australia) and the output in the non-traded sector is proxied by the unemployment rate since this is available monthly and a large portion of non-traded output comes from the labour-intensive service sector. The data are described in Appendix B. The VAR ordering reflects the dependent economy assumption that foreign prices are exogenous, but the exchange rate depends on these prices, non-traded output, and the money supply. The money supply is allowed to depend on

\textsuperscript{16}The Australian dollar floated in December 1983. The sample is constrained to start in January 1984 by the availability of the commodity price data. The VAR was tested with up to 6 lags resulting in a start date of July 1984.

\textsuperscript{17}The results for the VAR are consistent with the single equation approach used in Chapter 3 to estimate commodity currency models.
in absolute terms than the estimate for the Australian dollar.\textsuperscript{19} As noted earlier, the value of $\beta$ is sensitive to value of $\hat{i}_0$ around which the Taylor expansion is made, and so the precise magnitude of $\beta$ implied by the model must be treated with some caution.

Overall the model of biased expectations with regard to monetary shocks, but not commodity price shocks, appears to be consistent with the data. The predictions these assumptions have for a rapid response of the exchange rate to commodity price shocks, but delayed response to monetary shocks, are borne out by the VAR based on the dependent economy model of the exchange rate.

1.4.2 Peso Problems

The negative correlation of ex post expectation errors and the forward premium, which is implicit in the finding of $\beta < 0$, need not be caused by agents having systematically biased expectations, but rather may be a spurious small sample result. The expectations formed ex post by the researcher based on the whole sample could appear systematically biased if agents had to learn about a regime change, or a particular event occurred with an unrepresentative frequency. Notably if the learning was about a domestic monetary event, or the unrepresentative regime change concerned monetary policy, then exchange rate forwards would be biased while commodity futures would not.

The learning hypothesis, which is developed by Lewis (1988, 1995), initially appears to show promise as an explanation for the relative bias in exchange rates and commodity prices for Australia. In the early 1990s the monetary regime in Australia seemingly changed, as money growth and inflation declined. Further, this 'regime change' was not 'announced' until the first reference to inflation targeting by the central bank in 1993. If agents learned about this change slowly then their expectations over this period would, ex post, be recognised to have been systematically biased. But as Figure 1-6 shows, agents appear to have displayed rapid learning as the forward premium declined at the same time as the regime change indicated by the fall in money growth. Evans (1996) notes that if agents learn about a change rapidly, or the period of learning is a small portion of the sample, which is also the case in this sample, then

\textsuperscript{19}This value is similar in magnitude to the estimate for the more recent period, as shown in Figure 1-2, although the standard errors for estimates based on such short samples are wide.

41
the learning process will not cause the substantial ex post expectation errors that are needed to deliver the forward bias of exchange rates. Further, as Figure 1-2 showed, the $\beta$ from the exchange rate forward premium regression is negative even for subsamples excluding this period of 'learning'. This evidence implies the apparent change in monetary regime in the sample is not the explanation for the disparate relative degrees of bias.

The alternative peso problem explanation relates to anticipated regime changes. Evans and Lewis (1995) and Kaminsky (1993) demonstrate that exchange rates in the 1970s and 1980s can be classified as having followed a Markov switching process of appreciating and depreciating regimes. They suggest that agents' rational expectations about these regime changes leads to seemingly systematic expectation errors that display a spurious small sample correlation with the forward premium. The absence of such a spurious small sample correlation in commodity prices indicates that for this small sample correlation to be the cause of the bias of exchange rate forwards it would have to come from some other factor, likely domestic monetary policy. But the consistency of the finding that exchange rate forwards are biased, both across countries
and time, strongly suggests that expected changes in one country’s domestic monetary policy is unlikely to be the explanation of the puzzle.\textsuperscript{20} While peso problems may contribute to particular episodes of bias, the complete answer to the puzzle lies elsewhere.

1.4.3 Monetary policy

The observed forward bias of exchange rates, but near absence of bias for commodity prices, could potentially result from the practice of domestic monetary policy. McCallum (1994) proposed that monetary policy aimed at smoothing changes in exchange rates can lead to the observed bias in forward exchange rates.\textsuperscript{21} This section applies a monetary rule akin to McCallum’s, in which the central bank is assumed to dislike both sharp changes in policy and large changes in the exchange rate, to the dependent economy model from Section 1.3. The monetary policy rule, which summarises these goals and a potential dependence on the change in commodity prices, is given by

\[ \Delta m_t = A \Delta m_{t-1} - B \Delta s_t + C \Delta p^*_t + \zeta_t \] (1.22)

The persistence of monetary policy is governed by the parameter $A$. The use of monetary policy to counteract changes in the exchange rate is shown by the negative coefficient on exchange rate returns, $-B$.

While expectations are assumed to be rational, two possible deviations from the expected depreciation implied by equation (1.10) are considered. McCallum’s framework requires that there are random deviations from UIP in order for monetary policy to cause the forward bias, although he does not elaborate on their cause. These shocks are assumed to be persistent in order to generate the observed persistence in the forward premium. This deviation from the general equilibrium framework is also taken here in order to demonstrate the impact of monetary policy on the exchange rate. The setting used here also allows for the possibility that part of the one period change in commodity prices is predictable, as Section 1.1 showed is in

\textsuperscript{20}Estimation of the Markov switching model of exchange rates showed that when the sample is extended using more recent data the pattern of long swings is no longer present. While this suggests there are more than two regimes, it also diminishes the likelihood that agents rationally expect such changes in regime.

\textsuperscript{21}Meredith and Ma (2002) extend McCallum’s analysis to show that the central bank’s objective with regard to changes in the exchange rate can result from policy objectives with regard to output and inflation.
fact the case.

The expected depreciation then is given by

\[
E_t \Delta s_{t+1} = \left\{ \frac{\tilde{\delta}}{1 + \tilde{\delta}} \sum_{j=0}^{\infty} \left( \frac{1}{1 + \delta t} \right)^j E_t \Delta m_{t+1+j} \right\} + \varphi q_t - \eta_t \quad (1.23)
\]

where \( \eta_t = D\eta_{t-1} + u_t \) is the persistent deviation from UIP, and \( \varphi q_t \) is the discounted sum of expected commodity price changes. For simplicity it is assumed that there is no additional information regarding commodity price changes over a horizon of greater than one period so that the discounted sum of expected commodity price changes is a multiple of the expected one period change \( E_t \Delta p^*_{X,t+1} = q_t = \Delta p^*_{X,t+1} - \xi_{t+1} \).

Solving the model involves conjecturing a solution and then demonstrating it is consistent with the system given by (1.22) and (1.23). Changes in the exchange rate and money supply are proposed to be a function of the state variables, \( \zeta_t, \eta_{t-1}, u_t, q_{t-1} \) and \( \xi_t \):

\[
\Delta m_t = \phi_1 \zeta_t + \phi_2 \eta_{t-1} + \phi_3 u_t + \phi_4 q_{t-1} + \phi_5 \xi_t \quad (1.24)
\]

\[
\Delta s_t = \phi_6 \zeta_t + \phi_7 \eta_{t-1} + \phi_8 u_t + \phi_9 q_{t-1} + \phi_{10} \xi_t \quad (1.25)
\]

Substituting equations (1.24) and (1.25), and their expectations, into equations (1.22) and (1.23), and then equating coefficients on the state variables, the expressions for the changes in the money supply and exchange rate are found to be

\[
\Delta m_t = \frac{B\theta}{BD\tilde{\delta} - (A - D) \theta} (D\eta_{t-1} + u_t) \quad (1.26)
\]

\[
\Delta s_t = \frac{1}{B} \zeta_t + \frac{\theta}{BD\tilde{\delta} - (A - D) \theta} [(A - D) \eta_{t-1} - u_t] + \varphi (q_{t-1} + \xi_t) \quad (1.27)
\]

where \( \theta = 1 + \tilde{\delta} - D \). The reduced form solution for changes in the money supply is a function only of the deviations from UIP, notably it does not depend on commodity prices. Changes in the exchange rate by contrast reflect the shocks to UIP, the shocks to monetary policy and commodity price changes. The 'Fama coefficient' will depend on the covariance between the
interest differential and the realised change in the exchange rate and is calculated as

\[
\text{plim} \left( \hat{\beta} \right) = \frac{\text{cov} \left( i_t - i_t^*, \Delta s_{t+1} \right)}{\text{var} \left( i_t - i_t^* \right)}
\]

(1.28)

\[
= \frac{\text{cov} \left( E_t \Delta s_{t+1} + \eta_t, \Delta s_{t+1} \right)}{\text{var} \left( E_t \Delta s_{t+1} + \eta_t \right)}
\]

(1.29)

\[
= \frac{(1 + \tilde{\delta} - D)(A - D) + \Theta}{DB\tilde{\delta} + \Theta}
\]

(1.30)

where \( \Theta = \frac{\phi^2 (BD\tilde{\delta} - (A - D)(1 + \tilde{\delta} - D))^2}{BD\tilde{\delta}} \). If there is no variability in the predictable change in commodity prices, so \( \sigma_q^2 = 0 \), then this reduces to \( \text{plim} \left( \hat{\beta} \right) = \frac{(1 + \tilde{\delta} - D)(A - D)}{B} \). Assuming the UIP shocks are not explosive, \( D < 1 \), the first term is positive and so \( \beta \) will be negative if the shocks to UIP demonstrates greater persistence than monetary policy, \( D > A \).\(^{22}\) Because of the number of parameters and the endogenous relationship between policy and the exchange rate the model is not identified and so the parameter values can’t be estimated. Instead, the consistency of this framework with empirical observations can be gauged by calibrations. For example, if \( A = 0.85 \), \( D = 0.9 \) and \( B = 0.2 \) then the implied value for \( \beta \) is \(-0.75\), close to the empirical estimate.\(^{23}\) But depending on parameter values, \( \beta \) can take on any value, either positive or negative.

If there are predictable changes in commodity prices \( \sigma_q^2 > 0 \), and so \( \Theta > 0 \), the range of parameters for which \( \beta \) will be heavily downward biased is greatly reduced. The value of \( \beta \) will depend on the ratio of the variances of predictable commodity price changes to the UIP shocks, \( \sigma_q^2 / \sigma_\eta^2 \). While an estimate of the volatility of the commodity forward premium is given in Table 1.2, because the shocks to UIP are unobserved, and McCallum does not provide a foundation for their existence, their volatility cannot be measured. In the absence of an alternative assumption, suppose these variances are equal. Calibrations then indicate \( \beta \) cannot be negative.\(^{24}\) For \( \beta \) to

\(^{22}\) This is the opposite result to McCallum, who finds that the UIP shocks must demonstrate greater persistence than the policy function, \( D > A \), in order for \( \beta < 0 \). The difference arises because in the current set up the interest rate is a function of the discounted stream of expected future money growth rather than being the policy variable as in McCallum. In this setting the interest rate moves in the same direction as the change in money, which is being used to offset the UIP shock. Despite this technical difference, the framework still demonstrates McCallum's point that monetary policy can induce negative covariance between interest rates and the exchange rate.

\(^{23}\) The calibration uses \( \phi = -0.7 \), based on estimates in Chapter 3, and \( \tilde{\delta} = 0.05 \), as used in Section 1.4.1.

\(^{24}\) The minimum occurs for \( A = D \to 0 \) implying that \( \beta \to 0 \).
be as low as the $-0.75$ obtained above, the variance of the shocks to UIP must be of the order of six times that of the predictable commodity price changes.\footnote{This value of beta occurs for a range of values of the parameters $A$, $B$ and $D$. One such combination is $A = 0.58$, $D = 0.99$, $B = 0.20$.}

This simple model demonstrates that the practice of monetary policy could induce bias in exchange rate forwards despite its absence in commodity futures. But the model requires strong assumptions with little justification, such as the shocks to UIP. Further, calibrations demonstrate that if the unbiased predictability of commodity prices is taken into consideration, the variance of these shocks to UIP must be very large, substantially larger than the variance of the predictable component of commodity prices.

1.5 Conclusion

This paper observed that a portfolio of commodity futures does not exhibit the same bias as do exchange rate forwards. This may be of only passing interest if it were not for the fact that some exchange rates have a close relationship with commodity prices. The nature of this relationship is developed in the dependent economy model. In this model, in addition to being a function of the price of commodity exports, the exchange rate is also a determined by the price of imports, the non-traded output and the domestic money supply. The behaviour of, and expectations about, the money stock appears to be the most likely cause of the bias in exchange rate forwards despite the absence of bias in commodity futures. Three explanations for the exchange rate forward bias that depend on the domestic money stock were considered. Explanations in the peso problem class appear unlikely to be the cause of the puzzle. Learning about an in-sample regime change was rapid, while the existence of forward bias across exchange rates and sample periods suggests it is not sample specific, as would likely be the case with a regime change. Monetary policy aimed at smoothing the exchange rate could downward bias the estimate of beta, but calibrations suggest that if expectations about commodity futures are unbiased, the monetary policy response is unlikely to be able to cause the full extent of the bias. The explanation that depends on systematic bias in expectations about the monetary process appears consistent with the empirical facts, in particular the rapid response of exchange rates to commodity shocks but delayed response to monetary shocks. However this explanation requires
a very strong assumption, that agents are capable of forming rational expectations with regard
to commodity prices, but not with regard to the money process. Perhaps the slow adaptation of
expectations to monetary shocks stems from institutional factors or other economic variables,
but before this explanation can be treated with any confidence, greater evidence as to the cause
of biased monetary expectations must be uncovered.
Bibliography


[16] Freebairn, John (1990) "Is the $A a commodity currency?" in K. Clements and J. Freebairn (Eds.) Exchange Rates and Australian Commodity Exports, Centre for Policy Studies, Monash University and Economic Research Centre, The University of Western Australia, Melbourne, 6-30.


Appendix A – Solutions

Dependent economy model

Consumption

All exogenous variables in the model are log normally distributed. Further, their log levels are random walks – the expected future value is simply the current value.

The three main first order conditions from maximising equation (1.4) with respect to the budget constraint are

\[
\frac{C_{Tt}P_{Tt}}{1 - \gamma} = \frac{C_{Nt}P_{Nt}}{\gamma} \quad (1.31)
\]

\[
\frac{P_{Tt}}{C_{t}^{p}P_{t}} = \beta (1 + r) E_{t}^{m} \left\{ \frac{P_{Tt+1}}{C_{t+1}^{p}P_{t+1}} \right\} \quad (1.32)
\]

\[
\chi \left( \frac{M_{t}}{P_{t}} \right)^{-\delta} C_{t}^{p} = 1 - \beta E_{t}^{m} \left\{ \left( \frac{C_{t}}{C_{t+1}} \right)^{\rho} \frac{P_{t}}{P_{t+1}} \right\} \quad (1.33)
\]

All shocks are log normally distributed and so $C_{T}$ is assumed to be log normally distributed. Since $C_{N} = Y_{N}$ is also log normal, this will imply aggregate consumption, $C$, is also log normal. Domestic prices are also assumed to have log normal distributions. In the solution these assumptions hold. Using the log normality, equation (1.32) can be expanded as

\[
-\gamma E_{t}^{m} \{ p_{Tt+1} - p_{Tt} \} + \rho \gamma E_{t}^{m} \{ c_{Nt+1} - c_{Nt} \} + \rho (1 - \gamma) E_{t}^{m} \{ c_{Tt+1} - c_{Tt} \}
\]

\[
+ \gamma E_{t}^{m} \{ p_{Nt+1} - p_{Nt} \}
\]

\[
= \frac{1}{2} \text{var} (p_{Tt+1} - \rho c_{t+1} - p_{t+1}) + \ln [\beta (1 + r)]
\]

\[
= K_{1}
\]

Substituting in equation (1.31) this simplifies to equation (1.5) in the text. Since non-traded consumption must equal non-traded output, which has an exogenously given log random walk distribution, the expectation of traded consumption is given by

\[
E_{t}^{m} \{ c_{Tt+1} \} = c_{Tt} + \frac{K_{1}}{\rho (1 - \gamma) + \gamma}
\]
Then the expected value of traded good consumption is given by

\[ E_t^m \{ C_T \} = C_T \exp(K_3) \]

where \( K_3 = \frac{(s-t)K_1}{\rho(1-\gamma)+\gamma} + \frac{(s-t)}{2} \text{var} (c_{Tt+1} + p_{Tt+1}) \).

The intertemporal budget constraint, that the discounted present values of traded consumption and the exported good must be equal, then implies that

\[ \frac{1+r}{1+e^{\frac{1}{2}\sigma_X^2}} p_{Xt}^* = \frac{1+r}{1+e^{\frac{1}{2}\sigma_X^2}} C_T p_{Tt}^* \]

Assuming the parameter values satisfy \( K_3 = \frac{1}{2} \sigma_X^2 \), ensuring that a steady state solution exists, then tradeables consumption is indeed log normally distributed and follows a random walk governed by

\[ c_{Tt} = p_{Xt}^* - p_{Tt}^* \]

**Exchange rate**

The first order conditions with respect to holdings of the domestic nominal bond and money balances are

\[ \frac{1}{C_t^\rho P_t} = \beta (1+i_t) E_t^m \left\{ \frac{1}{C_{t+1}^\rho P_{t+1}} \right\} \]

\[ 1 - \chi \left( \frac{M_t}{P_t} \right)^{-\delta} C_t^\rho = \beta E_t^m \left\{ \left( \frac{C_t}{C_{t+1}} \right)^\rho \frac{P_t}{P_{t+1}} \right\} \]

Combining these gives

\[ \chi \left( \frac{M_t}{P_t} \right)^{-\delta} C_t^\rho = \frac{i_t}{1+i_t} \]

The non-stochastic steady state is defined by a constant rate of growth of money such that equation (1.36) is constant.

\[ \chi \left( \frac{M}{P} \right)^{-\delta} \bar{C}^\rho = \frac{\bar{i}}{1+i} \]
The left hand side of equation (1.35) can then be log linearised around the non-stochastic steady state, equation (1.37). Doing so gives equation (1.7) in the text. The right hand side of (1.35) is log linear, because all shocks are log normally distributed, and so does not need to be approximated. Taking the log of the right hand side, and equating it to the log approximation of the left hand side, equation (1.7), gives equation (1.8) in the text.

**Derivation of biased market expectations**

Agents' beliefs of the money process are given by

\[
m_t = z_t + \nu_t \quad \quad (1.38)
\]

\[
z_t = \lambda z_{t-1} + \epsilon
\]

where \( var(\nu) = \sigma^2_\nu \) and \( var(\epsilon) = \sigma^2_\epsilon \). The Kalman algorithm for updating forecasts, as used by Gourinchas and Tornell and described in Hamilton (1995), is

\[
E_t^m z_{t+1} = \lambda E_{t-1}^m z_t + \lambda P (P + \sigma^2_\nu)^{-1} (m_t - E_{t-1}^m m_t)
\]

(1.39)

The expectations of money are then given by

\[
E_t^m m_{t+1} = E_t^m z_{t+1}
\]

\[
= \lambda E_{t-1}^m z_t + \lambda P (P + \sigma^2_\nu)^{-1} (m_t - E_{t-1}^m m_t)
\]

\[
= \lambda E_{t-1}^m m_t + \lambda k (m_t - E_{t-1}^m m_t)
\]

\[
= k \lambda m_t + (1 - k) \lambda E_{t-1}^m m_t
\]

where \( k = P (P + \sigma^2_\nu)^{-1} \) and \( P \) is the mean-squared error of the expectation as given by

\[
P_{t+1} = \lambda^2 \left[ P_t - \sigma^2_\nu (P_t + \sigma^2_\nu)^{-1} \right] + \sigma^2_\epsilon
\]
As in Gourinchas and Tornell it is assumed that this estimate of the prediction error variance has converged to a stable value

\[
P = \lambda^2 \left[ P - P^2 (P + \sigma_v^2)^{-1} \right] + \sigma_e^2
\]

\[
P = \frac{-\left((1 - \lambda^2) \theta - 1\right) + \Theta}{2/\sigma_e^2}
\]

where \(\theta = \sigma_v^2/\sigma_e^2\) summarises the agents' belief of the noise-to-signal ratio in the money process, and \(\Theta = ((1 - \lambda^2) - 1) - 4\theta\). Since \(P > 0\) the constant \(k\) is between zero and one.
Appendix B – Data

Exchange Rate Data

Australian dollar spot and 3-month forward rates for the first business day of each month were obtained from Bloomberg (AUD Curncy, AD3M Curncy). The data are the average of buy and sell rates, however, Bekaert and Hodrick (1993) found this introduces no bias relative to the use of correctly aligned transaction prices. Since futures data often contain occasional measurement error, as shown by Maynard and Phillips (2001), the data were filtered mechanically and manually. For the days used in this study no obvious measurement errors were found in the data.

Commodity Data

The commodity futures indices are constructed using data from three sources. The weights are derived from the export-share based weights used in the Reserve Bank of Australia (RBA) commodity price index. The weights, listed in Table 1.4 are rescaled to account for the omission of several components for which futures or forwards prices are not available. The commodity price data come from the London metal exchange (LME) and the Futures Industry Association (FIA).

LME forwards

The data from the LME are daily spot and 3-month forward prices for aluminium, copper, nickel, zinc and lead. The FIA data, unlike the LME data, are forwards rather than futures. There is a substantial literature examining the theoretical and empirical differences between forwards and futures. Cox, Ingersoll and Ross (1981), show that forward and futures prices can differ if the interest rate is stochastic. Empirically it is found that there is no discernible difference between the prices of futures and forwards (viz. French (1983), and Cornell and Reinganum (1981)) and so there seems little problem with combining these series. While some of the historical data prior to 1989 had to be converted from British pounds to dollars – using the daily spot and 3-month forward rates – almost all of the LME data are for trades in US dollars. A three year overlap of the LME copper series which traded in pounds was compared with the
FIA copper series trading in dollars and found to have almost identical monthly movements, indicating the currency conversion did not introduce significant bias.

**FIA futures**

The FIA data come from numerous exchanges, as shown in Table 1.4, all in the USA and trading in US dollars. Cash prices for commodities can be for a different grade commodity, in a different location, or include discounts, and as such they are not necessarily a good measure of the spot price. For this reason, studies using commodity futures data frequently use expiring contracts as the spot price. Most commodity contracts are deliverable in the first 3 weeks of the contract month. Following Fama and French (1987) the price on the first day of the delivery month, when the contract can be delivered and so the contract price should have converged to the implicit spot price, but the contract is still liquid, is used as the spot price.

**Constructing observations for missing months** For most commodities there are not contracts expiring in each month. In order to construct the commodity index it is necessary to have monthly observations for each series. Following Pindyck (1993) log linear extrapolations of longer horizon futures are used to construct a spot price, or 3-month futures price, in months in which there is no observation. For example, if there is no 3-month futures contract in a given month, but there are 4- and 5-month contracts, these are used to construct an implied 3-month contract, adjusting for the number of days between contracts. Specifically, if \( F_{0t} \) is the futures price to be constructed, and \( F_{1t} \) and \( F_{2t} \) are the two longer contracts, the formula used is

\[
F_{0t} = F_{1t} \left( F_{1t}/F_{2t} \right)^{n_{01}/n_{12}}
\]  

(1.40)

where \( n_{ij} \) is the number of days between the expiration of contracts \( i \) and \( j \). The accuracy of this method of extracting futures prices was assessed by comparing the implied price with the observed price in the months where there was an actual observation. Table 1.5 shows that the series of actual and implied prices are highly correlated, and that the mean percentage error is close to zero. The mean percentage squared error is larger for agricultural commodities for which seasonal factors are likely to be relevant, and indicates that this extrapolation method introduces some measurement error. This is unavoidable. Fortunately, the similar behaviour
Table 1.4: Composition of Commodity Futures Index.

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<td></td>
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<tr>
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<td>3.8</td>
</tr>
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<tr>
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<td></td>
<td></td>
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<tr>
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<td></td>
</tr>
<tr>
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<tr>
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<tr>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Energy</td>
<td>LNG</td>
<td>0.9</td>
</tr>
<tr>
<td></td>
<td>Coking Coal</td>
<td>16.6</td>
</tr>
<tr>
<td></td>
<td>Steaming Coal</td>
<td>8.3</td>
</tr>
</tbody>
</table>

Proportion of RBA index: 100.0 52.6 12.2 40.4

* index – indicates the use of an index formed from the three Wheat contacts
of the futures series and the LME forwards series, for which extrapolations are not needed, suggests the measurement error does not unduly affect the results.

<table>
<thead>
<tr>
<th>Contract</th>
<th>Exchange</th>
<th>Mean % Error</th>
<th>RMS % Error</th>
<th>Correlation with actual series</th>
<th>Mean % Error</th>
<th>RMS % Error</th>
<th>Correlation with actual series</th>
</tr>
</thead>
<tbody>
<tr>
<td>Cattle - Feeder</td>
<td>CME</td>
<td>0.40</td>
<td>2.09</td>
<td>0.992</td>
<td>0.33</td>
<td>1.84</td>
<td>0.994</td>
</tr>
<tr>
<td>Cattle - Live</td>
<td>CME</td>
<td>0.11</td>
<td>3.78</td>
<td>0.968</td>
<td>-0.07</td>
<td>2.92</td>
<td>0.979</td>
</tr>
<tr>
<td>Wheat</td>
<td>KCBT</td>
<td>-0.06</td>
<td>2.74</td>
<td>0.991</td>
<td>0.47</td>
<td>2.62</td>
<td>0.989</td>
</tr>
<tr>
<td>Wheat - White</td>
<td>MINN</td>
<td>-0.53</td>
<td>2.70</td>
<td>0.987</td>
<td>0.46</td>
<td>2.41</td>
<td>0.89</td>
</tr>
<tr>
<td>Wheat</td>
<td>CBT</td>
<td>0.76</td>
<td>3.82</td>
<td>0.989</td>
<td>0.66</td>
<td>3.41</td>
<td>0.990</td>
</tr>
<tr>
<td>Rice - Rough</td>
<td>CBT</td>
<td>-0.68</td>
<td>4.82</td>
<td>0.983</td>
<td>-0.01</td>
<td>6.00</td>
<td>0.948</td>
</tr>
<tr>
<td>Corn</td>
<td>CBT</td>
<td>-0.03</td>
<td>3.19</td>
<td>0.984</td>
<td>-0.12</td>
<td>3.26</td>
<td>0.987</td>
</tr>
<tr>
<td>Sugar No. 11</td>
<td>CSCE</td>
<td>1.98</td>
<td>7.85</td>
<td>0.993</td>
<td>0.22</td>
<td>2.83</td>
<td>0.999</td>
</tr>
<tr>
<td>Cotton No. 2</td>
<td>NYCE</td>
<td>0.09</td>
<td>6.69</td>
<td>0.937</td>
<td>0.04</td>
<td>3.83</td>
<td>0.981</td>
</tr>
<tr>
<td>Gold</td>
<td>NYNEX</td>
<td>-0.09</td>
<td>0.27</td>
<td>1.000</td>
<td>0.00</td>
<td>0.08</td>
<td>1.000</td>
</tr>
<tr>
<td>Copper - HG</td>
<td>NYNEX</td>
<td>0.19</td>
<td>0.87</td>
<td>0.999</td>
<td>-0.04</td>
<td>0.60</td>
<td>0.999</td>
</tr>
</tbody>
</table>

**Substitute contracts** For several commodities there are no futures markets, or were no active futures markets for the full length of the sample. Where possible contracts for substitute commodities (e.g. for rice) or complement commodities (e.g. for the metals) were used instead, as shown in Table 1.4. In each case the constructed or spliced series was compared to spot price data used in the RBA commodity price index to ensure they accurately represent movements in the given commodity. The correlations with the RBA components are shown in Table 1.4. The constructed index of futures covers just over half of the weights in the RBA commodity index.

**Other data**

**Money** Financial reform in Australia over the 1980s and 1990s resulted in an increasing proportion of funds being captured by narrow monetary aggregates. To minimise the influence of this measurement issue the broadest monetary measure, broad money (BM), is used. The level of BM is taken from RBA Bulletin table D3 and backcast using monthly growth rates from table D1 which account for the largest of breaks to the series.

**Import price** A monthly import price series is not available for Australia to use as the tradable good price series. Instead the export price from industrial countries from the IMF IFS
database (line 11074..DZF) is used.

**Unemployment rate** The unemployment rate, taken from RBA Bulletin table G6, is used to proxy non-traded output.
Chapter 2

Identifying the efficacy of central bank intervention

2.1 Introduction

Since the introduction of floating exchange rates the use, and efficacy, of intervention in the foreign exchange market has been a controversial topic.\footnote{This chapter is co-authored with Roberto Rigobon} Most central banks have at times engaged in frequent intervention, and at other times followed a more laissez faire approach to the exchange rate.\footnote{Schwartz (2000) suggests that intervention is a dying practice despite the continued use of active intervention by the ECB and Bank of Japan. However, the 18 central banks that responded to a survey reported in Neely (2001) believe it affects the exchange rate. Traders' survey responses in Cheung and Wong (2000) indicate they also believe intervention has an effect on exchange rates.} No doubt the observed disparate range of policies between central banks, and within individual central banks over time, can in part be attributed to the lack of accord on the effectiveness and consequences of central bank intervention. Two key questions remain unresolved: how effective is foreign exchange intervention, and, if it is effective, through which channel does it act?

A critical barrier to answering these questions has been overcoming the endogeneity of changes in the exchange rate and intervention. The central hypothesis is that intervention changes the exchange rate. But at the same time, the decision to intervene is not independent of the movements in the exchange rate. Moreover, even once a central bank has decided to
intervene, the quantity of currency it buys or sells will typically depend on the response of the exchange rate to its trades.

The literature has typically dealt with the simultaneous equations problem by assuming that the contemporaneous decision of the central bank is independent of the current innovations to the exchange rate. This is a strong assumption. For example, it assumes that the central bank does not change its selling or buying behaviour by assessing the impact its actions have had on the exchange rate. On the other hand, there is strong evidence in stock markets that big players act strategically when they are unwinding large positions. Therefore, why should we expect the same behaviour is not optimal for a central bank?

In this paper we use an alternative identification method to solve the problem of simultaneous equations. We use daily Reserve Bank of Australia (RBA) interventions data over the period 1986 to 1993, which contains a dramatic change in intervention policy that we use for identification. We show that the estimates we obtain have the correct sign and are significantly larger than those found with more standard methods. This is exactly the direction we would have expected if endogenous variables is an important source of the bias. Further, the vast majority of the effect of an intervention on the exchange rate is found to occur during the day in which it is conducted with a smaller impact on subsequent days. This explains why small effects are usually found when lag values are used in the typical OLS specifications. The major contribution of this paper is to provide some evidence on the contemporaneous effectiveness of intervention. Although, our methodology does not indicate the channels through which intervention operates, it provides an improvement from previous estimates obtained in the literature.

We concentrate our analysis on sterilised interventions, but we do not distinguish between secret and public interventions. While undoubtedly this is an important distinction, most and in particular the largest, interventions are public. Certainly, future research should reconsider this issue. In this paper the focus of our attention is the estimation problem. Indeed, we think that the simultaneous equations problem is the crucial aspect limiting our understanding of how effective policy is.

The identification assumption we use is based on the fact that in 1991 the RBA decided to change their policy regarding foreign exchange rate interventions. We interpret this shift in
policy as exogenous, which is an important ingredient in our solution to the problem. Other countries have also changed "exogenously" their policies, but typically central banks endogenously respond to the conditions in the market. Again, future research should endogenise the policy decision and extend the present analysis to deal with a more general framework. However, as is argued in the text below it is the case that the Australian central bank decision to change their method of intervention is unrelated to other macro events.

This chapter proceeds as follows. There is a brief review, in Section 2.2, of central bank intervention practices and the associated literature. A description of the data and discussion of RBA intervention follow this, in Section 2.3. Section 2.4 outlines the identification and estimation methodology used in this study. The results are presented in Section 2.5, followed by conclusions.

2.2 Review of the literature on Central bank intervention.

This section briefly reviews the literature on central bank intervention, focussing on the simultaneous relationship between, and temporal behaviour of, exchange rate returns and intervention. For more general and extensive reviews see Sarno and Taylor (2001), Dominguez and Frankel (1993b) and Edison (1993).

Empirical studies, and statements by central banks, suggest that central banks intervene in foreign exchange markets to slow or correct excessive trends in the exchange rate, i.e. they "lean against the wind", and to calm disorderly markets (for example Lewis (1995b) and Baillie and Osterberg (1997b)). The survey responses of central banks in Neely (2001) suggest that these factors continue to drive the decision to intervene. A recent study for Australia by Kim and Sheen (2000) has similar conclusions. Importantly, when central banks intervene they trade in blocks throughout the day. As Neely (2001) reports, their subsequent trades are conditional on the response of the exchange rate to their earlier trades.

Two main channels have been suggested through which sterilised intervention can affect the level of the exchange rate: the portfolio balance channel and the signalling channel. Interven-

---

3Other reasons occasionally cited by central banks include to target particular exchange rates, or to support other central banks. We do not explicitly consider the impact of intervention on conditional exchange rate volatility, see Rogers and Siklos (2001) for Australia or more generally Dominguez (1998) and Bonser-Neal and Tanner (1996).
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soon as traders digest the information contained within the intervention. Goodhart and Hesse (1993), Peiers (1997), Dominguez (1999) and Chang and Taylor (1998) find that interventions that are intended to be visible are typically reported by news services within 10 minutes to 2 hours, by which time it is often no longer ‘news’ to traders. Any effect through the portfolio balance channel is also likely to be rapid as bond holders quickly respond to the change in the relative supplies in the highly liquid market for government securities. Indeed, Neely (2001) reports that the majority of central banks believe the full effects of intervention are reflected in the exchange rate within a matter of hours.

Despite the evidence of the rapid response of the exchange rate to intervention, studies using daily data have often abstracted from the endogeneity of intervention and exchange rate determination by only including lagged intervention (for example Baillie and Osterberg (1997a) and Lewis (1995b)). While intervention may still have an effect on the days subsequent to the initial trades, omitting the contemporaneous intervention prevents measurement of the immediate impact and is likely to bias other coefficient estimates. Other studies that include contemporaneous intervention, such as Kaminsky and Lewis (1996) and Kim et al (2001), typically obtain an incorrectly signed contemporaneous coefficient, suggesting that purchases of the domestic currency cause it to depreciate. Seemingly what they capture is the policy function coefficient that represents central banks’ tendency to “lean against the wind”. The insignificant and incorrectly signed coefficients in many previous studies indicate that an accurate estimation of the impact of intervention on the exchange rate must incorporate the contemporaneous effect, and account for the endogeneity between these variables.

2.3 Reserve Bank of Australia Intervention Data

The estimation in this study uses daily interventions in the foreign exchange market by the RBA over the period from July 1986 to November 1993. This period is chosen because it contains a single distinct change in intervention policy, in October 1991, that allows us to identify the parameters in our exchange rate and intervention system. Additionally, the characteristics of intervention in this sample facilitate accurate estimation. Over the 7 years of data (1,930 daily observations) intervention is frequent (48 per cent of days) and often large (up to $A1.3 billion).
Another advantage of our sample is that other central banks do not intervene in the USD/AUD market and so we need only focus on the actions of one central bank.

The data are the daily net purchases of Australian dollars by the RBA and include all transactions made by the RBA, including those on behalf of the Government. Since the RBA can, and does, meet the Government’s demand for foreign currency transactions from its own reserves, it has discretion as to the timing of transactions on behalf of the Government. For this reason the inability to exclude transactions on behalf of the Government does not appear to be problematic. Similarly, Neely (1998) argues that not excluding client transactions for the Fed has little influence. Over the sample used, almost all RBA intervention were conducted in the spot market versus the US dollar (Andrew and Broadbent (1994)).\(^8\) Rankin (1998) reports that the RBA has always sterilised its interventions.\(^9\)

The exchange rate is measured as the US dollar value of an Australian dollar (USD/AUD). The intervention and exchange rate returns data are aligned to cover exactly the same 24-hour period, commencing at 9am Sydney time. This avoids inaccurate results from using misaligned data, which have potentially hampered previous studies on central bank intervention. Aligning the intervention and exchange rate returns data is important since central banks can, and do, intervene outside their business hours, as shown in Dominguez (1999), and stated for the RBA by Rankin (1998).\(^10\) The 24-hour period over which the data are measured will be referred to as a day. The intervention data for any given day will include transactions conducted in any of the major markets around the world on that calendar day.\(^11\)

Rankin (1998) outlines five distinct periods of intervention policy used by the RBA since the float of the Australian dollar in December 1983. Table 2.1 summarises the pattern of intervention over these five episodes. The two periods used in this study are the second and

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\(^8\)Over the past decade the RBA has moved to using a combination of spot market transactions and currency swaps, which together replicate a forward, to intervene and sterilise their intervention (see Rankin (1998)).

\(^9\)Dominguez and Frankel (1993b) and the references contained therein, suggest the Fed, Bundesbank and Bank of Japan have typically only partially sterilised their interventions.

\(^10\)Given Australian trading hours do not overlap with either London or New York trading hours, and BIS (1993) shows one-quarter of Australian dollar trading occurs in markets outside of Australian trading hours, this is potentially important.

\(^11\)For a few 24 hour periods that end on an Australian public holiday we do not have an observation in the Sydney 9am exchange rate series. Instead the Australian dollar exchange rate measured at noon EST (2am or 4am Sydney time, depending on daylight saving) by the New York Fed is used.
Table 2.1: Summary Statistics of Interventions.

<table>
<thead>
<tr>
<th>Regime</th>
<th>Full</th>
<th>I</th>
<th>II</th>
<th>III</th>
<th>IV</th>
<th>V</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Dec-83</td>
<td>Dec-83</td>
<td>Jul-86</td>
<td>Oct-91</td>
<td>Dec-93</td>
<td>Jul-95</td>
</tr>
<tr>
<td></td>
<td>Jul-00</td>
<td>Jun-86</td>
<td>Sep-91</td>
<td>Nov-93</td>
<td>Jun-95</td>
<td>Jul-00</td>
</tr>
<tr>
<td>Number of interventions</td>
<td>Sales</td>
<td>Purchases</td>
<td>Total number of days</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>1283</td>
<td>504</td>
<td>4219</td>
<td>99</td>
<td>780</td>
<td>15</td>
</tr>
<tr>
<td></td>
<td>780</td>
<td>143</td>
<td>1338</td>
<td>156</td>
<td>116</td>
<td>0</td>
</tr>
<tr>
<td>Probability of intervention</td>
<td>Unconditional</td>
<td></td>
<td>Conditional on $INT_{t-1} \neq 0$</td>
<td></td>
<td>Conditional on $INT_{t-1} = 0$</td>
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<tr>
<td></td>
<td>0.42</td>
<td>0.73</td>
<td>0.20</td>
<td>0.51</td>
<td>0.62</td>
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<tr>
<td></td>
<td></td>
<td>0.69</td>
<td>0.80</td>
<td>0.40</td>
<td>0.40</td>
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<td></td>
<td>0.69</td>
<td>0.80</td>
<td>0.40</td>
<td>0.40</td>
<td>0.24</td>
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<tr>
<td>Average intervention ($Am)</td>
<td>Sales</td>
<td>Purchases</td>
<td>Absolute value</td>
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<tr>
<td></td>
<td>46</td>
<td>81</td>
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<td>285</td>
<td>46</td>
<td>285</td>
<td>159</td>
<td>46</td>
</tr>
<tr>
<td>Maximum ($Am)</td>
<td>Sale</td>
<td>661</td>
<td>661</td>
<td>661</td>
<td>150</td>
<td>0</td>
</tr>
<tr>
<td></td>
<td>1,585</td>
<td>1,026</td>
<td>1,305</td>
<td>90</td>
<td>1,305</td>
<td>0</td>
</tr>
<tr>
<td></td>
<td>Purchase</td>
<td>20,404</td>
<td>16,206</td>
<td>19,853</td>
<td>22,608</td>
<td>28,995</td>
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<td>Money base</td>
<td>60,976</td>
<td>37,261</td>
<td>57,435</td>
<td>74,242</td>
<td>102,921</td>
</tr>
<tr>
<td>Memo items ($Am)</td>
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<td></td>
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</tr>
<tr>
<td></td>
<td>Low bound regime (Jul-86 to Sep-91)</td>
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<td></td>
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</tr>
<tr>
<td></td>
<td>High bound regime (Oct-91 to Nov-93)</td>
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<td></td>
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<td></td>
</tr>
</tbody>
</table>

Note: Purchases are of Australian dollars, sample starts 13 December 1983.
Source: Authors' calculations.

Figure 2-1: Proportion of days with intervention.
third. During the period immediately following the float of the Australian dollar, December 1983 to June 1986, the RBA allowed the dollar to float freely with only very small interventions. Our study commences using data from July 1986 when the RBA took on a distinctly proactive policy to foreign exchange market intervention. Interventions became more frequent (the RBA was active on 69 per cent of days) and substantially larger (the average absolute intervention was $A63 million). The change in policy in our sample occurs in October 1991 when the RBA all but ceased to make very small interventions. Large interventions continued to be used as in the preceding era as seen in Figure 2-1. As a result, the frequency of interventions was drastically reduced (to 24 per cent of days), and the average size of interventions increased substantially (to $A144 million). We do not consider the last two episodes when the RBA did not intervene from December 1995 to June 1995, or when interventions resumed in July 1995.

Table 2.2, drawn largely from Neely (2001), compares the frequency and size of RBA interventions with those of the Fed, the Bundesbank, and the Swiss National Bank. Over the period covered in the table, from the early 1980s to the end of the 1990s, the RBA intervened on 42 per cent of days, compared to between 4 and 12.5 per cent for the other central banks. Typical central bank interventions are tiny relative to the huge daily turnover in foreign exchange markets. The daily turnover in the USD/AUD market was estimated to be $US17.9 billion in 1992 by BIS (1993), although only $US4.8 billion of this involved a non-financial counterparty.\textsuperscript{12} While the Australian economy is substantially smaller than the US and German economies, and the Australian dollar less heavily traded than all three other currencies in Table 2.2, the RBA interventions are of a comparable size to those of the larger central banks.\textsuperscript{13} The large magnitude of some RBA interventions is also confirmed by comparing the maximum interventions to the money base and M1 in Table 2.1. In the two episodes considered in this study, the largest intervention is seen to be over 5 per cent of the money base, or 2 per cent of M1. Note that while the turnover in the foreign exchange market has increased substantially, and certainly did so through our sample, neither of the hypothesised transmission channels are affected by the

\textsuperscript{12} The average (absolute) intervention for the regime including 1992 was just 0.6 per cent of the total average daily turnover.

\textsuperscript{13} The BIS survey lists the USD/AUD currency pair as the 9th most traded. The average daily turnover for the USD/DEM, USD/JPY and USD/CHF were $US192.2b, $US154.8b, and $US48.8b. Only considering trades with non-financial counterparties the USD/AUD was the fifth most traded currency pair with the three pairs mentioned before respectively $US26.9b, $US32.3b, $US7.0b.
volume of turnover.

Table 2.2: Comparison of Interventions by Central Banks.

<table>
<thead>
<tr>
<th></th>
<th>Beginning</th>
<th>End</th>
<th>US dollar purchases ($USm)</th>
<th>Proportion of days with intervention (%)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td></td>
<td>min</td>
<td>max</td>
</tr>
<tr>
<td>US</td>
<td>DEM/USD</td>
<td>7/1/83</td>
<td>12/31/98</td>
<td>-797</td>
</tr>
<tr>
<td>JPY/USD</td>
<td>7/1/83</td>
<td>12/30/98</td>
<td>-951</td>
<td>722</td>
</tr>
<tr>
<td>Germany</td>
<td>DEM/USD</td>
<td>7/1/83</td>
<td>12/31/98</td>
<td>-833</td>
</tr>
<tr>
<td>Switzerland</td>
<td>CHF/USD</td>
<td>1/3/86</td>
<td>12/29/99</td>
<td>-545</td>
</tr>
<tr>
<td>Australia</td>
<td>AUD/USD</td>
<td>12/13/83</td>
<td>7/30/99</td>
<td>-932</td>
</tr>
</tbody>
</table>

Source: Australia is from authors’ calculations. All others from Neely (2001).

2.4 Identification of the effectiveness of foreign exchange rate intervention

In this section we use a simple model of central bank intervention to highlight the limitations of the standard methods and to show how we solve the identification problem.

The manner in which intervention affects the exchange rate can be seen from a generalised uncovered interest parity (UIP) relationship

\[ e_t = E\{e_{t+1}|\Omega_t\} + i_t - i_t^* + \eta_t \]  \hspace{1cm} (2.1)

where \( e_t \) is the exchange rate (value of domestic currency), \( \eta_t \) represents possible predictable deviations from UIP, and \( \Omega_t \) is the time \( t \) information set, which includes contemporaneous and past interventions, \( \{INT_j\}_{j=t}^{\infty} \subset \Omega_t \). Substituting forward this relationship, and suppressing other elements of the information set, the exchange rate is

\[ e_t = E \left\{ \sum_{j=0}^{T-1} [i_{t+j} - i_{t+j}^* + \eta_{t+j}] \right\} INT_t + e_{t+T} \]  \hspace{1cm} (2.2)

where \( e_{t+T} \) is the exchange rate at some distant point in the future. Under the portfolio balance channel, the relative supplies of domestic and foreign bonds changes so that the risk premium on domestic assets, which is part of \( \eta_t \), changes. This is then seen to affect the current
exchange rate. The signalling channel suggests that interventions are a commitment to future monetary policy and so represent a change in the expected future interest rate, \( E\{i_{t+j}|\text{INT}_t\} \), so changing the current exchange rate. Under either channel the exchange rate is a function of contemporaneous intervention, \( e_t = e_t(\text{INT}_t) \).

As noted earlier, central banks typically state that they intervene to slow or correct excessive trends in the exchange rate and to calm disorderly markets. Indeed, the governor of the RBA states that the RBA has used intervention in "circumstances where market imperfections are resulting in overshooting" and that "intervention can play a useful role in limiting extreme movements in the exchange rate" (Macfarlane (1998)). Since the RBA sterilises interventions, and is acknowledged to allow the Australian dollar to float quite freely, it seems reasonable to assume that intervention is not used as a separate policy tool with independent goals. Rather, intervention is focused on exchange rate outcomes. A simple representation of this policy, as used in Almekinders and Eijffinger (1996), is that the central bank's preferred level of intervention, or shadow intervention, \( \text{INT}^* \), would minimise squared deviations of the exchange rate from a moving target.

\[
L = (e_t(\text{INT}_t^*) - \bar{e}_t)^2
\]

(2.3)

Given the central bank is allowing the exchange rate to float, but doesn't want it to move 'too quickly', the target is taken to be a moving average of past values of the exchange rate, \( \bar{e}_t = \frac{1}{n}\sum_{j=1}^{n} e_{t-j} \). The optimal level of intervention will then by given by

\[
e_t(\text{INT}_t^*) - \bar{e}_t = 0
\]

(2.4)

\[
\Delta e_t(\text{INT}_t^*) = -\sum_{j=1}^{n-1} \left( \frac{n-j}{n} \right) \Delta e_{t-j}
\]

(2.5)

However, central banks do not intervene on every day, and very small interventions are extremely rare. Presumably there are some costs to intervention, possibly because the strength of signals is reduced if they are used too frequently. As a result the central bank only intervenes if the loss function would exceed some benchmark, or equivalently if the shadow intervention exceeds a given threshold, otherwise remaining absent from the market. Actual intervention
can then be represented as

\[ INT_t = \mathcal{I} \left( |INT_t^*| > \bar{INT} \right) \cdot INT_t^* \]  

(2.6)

where \( \mathcal{I}() \) is the indicator function.

Equations (2.2), (2.5) and (2.6) constitute a system that determine the exchange rate and intervention.

2.4.1 Setup of the estimation system

We generalise the previous framework to include unobservable variables that affect the exchange rate and the central bank's decision for intervention. The reason behind this extension is that we believe that there are factors that are unobservable at daily frequencies that have impact on both variables, such as liquidity shocks, macro shocks, etc. The model used is:

\[ \Delta e_t = \alpha INT_t + \gamma z_t + \varepsilon_t \]  

(2.7)

\[ INT_t = \mathcal{I} \left( |INT_t^*| > \bar{INT} \right) \cdot INT_t^* \]  

(2.8)

\[ INT_t^* = \beta \Delta e_t + z_t + \eta_t \]  

(2.9)

\( \Delta e_t \) is the observed exchange rate return at date \( t \) (a positive value is an appreciation), \( INT_t \) is the observed intervention (positive values are purchases of the domestic currency), and \( INT_t^* \) is a shadow intervention. The estimation procedure includes constants and potentially lags, but this simple version is sufficient to demonstrate the endogeneity problem.

Equation (2.7) is the reaction of the exchange rate to the central bank intervention. We assume that the exchange rate is affected by two types of shocks: \( \varepsilon_t \) which is a pure idiosyncratic shock to the exchange rate, which we assume has no direct impact on the intervention decision, and a common shock (\( z_t \)) which is assumed to move both the exchange rate and the central bank intervention decision. We explore the interpretation of these shocks below.

Equation (2.8) is the decision of the central bank to intervene or not. We assume that this decision is made entirely based on the shadow intervention. In other words, if the required intervention is large (larger in absolute terms than some threshold \( \bar{INT} \)) then the central bank
participates in the market, otherwise, it remains absent. Observe that implicitly we assume
that if the shadow intervention is larger than the threshold then the central bank intervenes,
and its intervention is exactly the shadow one.

Equation (2.9) determines the shadow intervention. We assume that it is affected by the
movements in the exchange rate, by the aggregate or common shock, and by some idiosyncratic
shock reflecting innovations to exchange rate policy. If the central bank aims to offset changes
in the exchange rate, i.e. lean against the wind, then $\beta$ will be negative. Equations (2.8) and
(2.9) together constitute the central bank’s reaction function; where the former reflects the
decision to intervene, and the later one determines the quantity, or size of intervention.

The policy shock ($\eta_t$) is interpreted as innovations in the exchange rate target that are
independent of the nominal exchange rate shocks ($\varepsilon_t$) and the common shock ($z_t$). The idea
is to separate idiosyncratic shocks to policy (such as trades on behalf of the government or
unwinding of positions) and to the exchange rate (for example, economic fundamentals) from
those shocks that we might expect to affect both variables (such as herding, liquidity, or shocks
to the exchange rate during periods of high conditional volatility). These common shocks
will affect how intervention takes place, and the exchange rate at the same time. We assume
that these shocks are i.i.d., with mean zero and variances $\sigma_\varepsilon$, $\sigma_\eta$, and $\sigma_z$. For simplicity in
the exposition we have assumed that all the variables have zero mean, but in the empirical
implementation it is important to include constants.

Finally, the parameter of interest is $\alpha$. If central bank intervention is effective, then pur-
chases of the domestic currency will appreciate the currency and so $\alpha$ will be positive.

The intuition for this model is that the central bank leans against the wind, so $\beta < 0$, either
to slow deviations from trend, or to calm volatile markets. Small changes are tolerable and
so the central bank does not bother intervening. On the other hand, if exchange rate returns
would otherwise be large, larger interventions would be required to counteract these and will
cause the central bank to enter the market. The shadow intervention ($INT^*_t$) summarises the
expected intervention if the central bank were to trade continuously in the foreign exchange
rate market.

This simple framework captures the two sources of simultaneity that exist in the data.
The first one is the endogenous decision of participation. The second one is the size of the
intervention and the change in the exchange rate once the decision of participation has been made. While the first source of bias has been widely acknowledged in the literature, the second has received very little attention. This is understandable. Finding instruments for the first one is hard, but some might be available. For the second one, this is much more difficult.

In this model there does not exist an instrument that can be used to solve the problem of simultaneous equations. More importantly, this bias is likely to be negative, pushing the estimate of $\alpha$ in equation (2.7) downward, possibly even negative, explaining most of the results found in the data.

It is important to mention that there are several aspects of central bank intervention that have been oversimplified in this model. First, there is no distinction between public and secret interventions. As was mentioned before, this has received considerable attention in the literature. In this paper we focus on the estimation problem. Second, we do not attempt to distinguish between sterilised and unsterilised interventions as the RBA states that all of its interventions are sterilised.

2.4.2 Identification through changes in intervention policy.

The problem of identification is easily shown by counting the number of unknowns, and the number of series we can measure, in the model. Under the assumption that we only observe the exchange rate, the size of the intervention, and its timing, then, aside from the means, we can compute only five moments from the data: the probability (or frequency) of intervention, the variance of the exchange rate when there is no intervention, and the covariance matrix when an intervention has taken place. However, in the model there are seven unknown coefficients that explain the behaviour of such variables: the parameters of interest ($\alpha$, $\beta$, and $\gamma$), the threshold of intervention ($INT$) and the three variances ($\sigma_z$, $\sigma_\eta$, and $\sigma_\xi$).\footnote{The estimation of means adds the same number of equations and unknowns to the system, thus, the problem of underidentification remains the same.}

The standard procedures in the literature use the following assumptions. First, that there are good instruments for the participation decision. Second, that either $\beta = 0$ or $\alpha = 0$ (exclusion restrictions). And third, that the instrument is correlated with $\eta_t$ but not with $z_t$. This set of assumptions seems rather strong. Central banks no doubt intervene based on their
most recent information set, which includes the change in the exchange rate during the day. Further, the fact that central banks know they have market power, which is the whole rationale why they think it is worth intervening in the first place, collides with the assumption that $\beta$ is zero. Central banks should be, and indeed are, strategic in their interventions. The alternative identification assumption that $\alpha = 0$ is similarly problematic in that it implies that central bank interventions don't have any effect on the day during which they are conducted. This contradicts significant circumstantial evidence.

The main contribution of this paper, is to relax these set of assumptions and use an alternative identification method that can deal with some of the econometric issues at hand. Obviously, we depend on another set of assumptions. We think those are weaker, in the sense that most of them are already imposed in the standard literature. But this is certainly the first pass at the problem using these alternative methods and further research should extend the present procedure. We discuss the caveats in detail at the end.

Our identification procedure is quite simple; in September 1991, the RBA changed its foreign exchange rate intervention policy. Following the change, the RBA all but ceased to conduct small interventions but continued to undertake larger interventions as before the change. In the model this would be summarised by a shift in $\overline{INT}$. Effectively, this means that there are two regimes. Under the assumption that the parameters and the variance of the shocks remain the same across both regimes, we have only eight unknowns (one more than before because we have two thresholds) but at least ten moments in the data.

Specifically, the basic model we estimate is the following

$$
\Delta c_t = c_e + \alpha INT_t + \gamma z_t + \varepsilon_t
$$

$$
INT_t^* = c_{INT} + \beta \Delta c_t + z_t + \eta_t
$$

$$
INT_t = \begin{cases} 
\mathbb{I} \left( |INT_t^* - c_{INT}| > \overline{INT}_t \right) \cdot INT_t^* & t < \hat{t} \\
\mathbb{I} \left( |INT_t^* - c_{INT}| > \overline{INT}_{\hat{t}} \right) \cdot INT_t^* & t > \hat{t}
\end{cases}
$$

(2.10)

Note that in this setup we allow for constants in the mean equations and below we extend the model to also include lags.\(^{15}\)

\(^{15}\)The lag structure takes into account that the exchange rate can only depend on observable variables. Thus, the lag of intervention used is $INT_{t-1}$ and not $INT_{t-1}^*$. However, for the shadow intervention equation we allow it to depend on the lag shadow realisation.
We estimate this model with, and without, lags and present both results. When the model is estimated without lags, there are eleven parameters of interest

\[ \{ c_e, c_{INT}, \alpha, \beta, \gamma, \pi, \overline{INT}_l, \overline{INT}_h, \sigma_e, \sigma_\eta, \sigma_\epsilon \} . \]

We compute the following moments in each of the regimes: the proportion of days with intervention; the mean exchange rate return; the mean intervention; the variance of the exchange rate on days with no intervention; and, the variance-covariance matrix on days with intervention. Furthermore, we compute the moments related to the serial correlation of the exchange rate, as well as the probability of consecutive interventions. This gives us a total of 20 moments, greater than the number of parameters, leaving our system over-identified.

We use simulated GMM to estimate the model. The general idea of this procedure can be easily understood by analyzing how other techniques estimate the coefficients. For example, when we use Maximum Likelihood the goal is to estimate the parameters using the mean and the variances. GMM extends that procedure and uses other moments. Simulated GMM is a further generalisation, in which we choose different moments and characteristics from the data, and "create" our own data using our auxiliary model to match those "moments". Indeed, all three techniques use auxiliary models for their estimation. On the one hand, in ML we use multinomial distributions, described only by means and variances. On the other hand, simulated GMM creates its own data within a well specified model to produce the statistics that we are interested in matching from the population.

In summary, the procedure is as follows: First, we create random draws of 20,000 observations for three uncorrelated shocks with unitary variance. The same set of shock variables are retained for the entire estimation procedure. Second, we simulate the model given some initial conditions. Third, we compute the moments in the simulated data and compare them with the sample moments. Finally, we iterate this procedure to search for the coefficients that minimise the distance between the population moments and the simulated ones.

To calculate the standard errors of the estimates we use the asymptotic distribution of the sample moments. Using the data, we bootstrap the exchange rate, intervention, and probabilities of interventions to produce a sequence of moments (100 of them). Then we estimate
the coefficients for each draw of the moments, computing the distribution of our coefficients. Because it is likely that the data is serially correlated, the bootstrap takes this into account.

2.5 Results

Before we present our results, it is illustrating to demonstrate the problems in estimation that arise from using more traditional methods. As noted earlier, frequently, researchers exclude the contemporaneous impact of intervention in an attempt to overcome the endogeneity. To account for the properties of exchange rate data, a GARCH error structure is typically used. Table 2.3 presents the results from an EGARCH(1,1) model of log changes in the USD/AUD exchange rate over the full sample, and in each of the high and low bound regimes. To highlight the impact of the simultaneity bias, we also present regressions that include the contemporaneous intervention. The coefficients on lagged intervention are uniformly small (around 0.005), and frequently the wrong sign, suggesting that RBA purchases of Australian dollars depreciated the exchange rate.

<table>
<thead>
<tr>
<th>Table 2.3: EGARCH(1,1) model of Interventions.</th>
</tr>
</thead>
<tbody>
<tr>
<td>Full sample (7/86 - 11/93)</td>
</tr>
<tr>
<td>constant</td>
</tr>
<tr>
<td>(0.000)</td>
</tr>
<tr>
<td>INT_t</td>
</tr>
<tr>
<td>(0.001)</td>
</tr>
<tr>
<td>INT_{t-1}</td>
</tr>
<tr>
<td>(0.001)</td>
</tr>
<tr>
<td>INT_{t-2}</td>
</tr>
<tr>
<td>(0.001)</td>
</tr>
<tr>
<td>constant</td>
</tr>
<tr>
<td>(0.115)</td>
</tr>
<tr>
<td>(\varepsilon_t / \sigma_{t-1})</td>
</tr>
<tr>
<td>(0.025)</td>
</tr>
<tr>
<td>(\varepsilon_{t-1} / \sigma_{t-1})</td>
</tr>
<tr>
<td>(0.014)</td>
</tr>
<tr>
<td>(\log(\sigma^2_{t-1}))</td>
</tr>
<tr>
<td>(0.010)</td>
</tr>
</tbody>
</table>

Model:

\[
\Delta e_t = c + \phi_0 INT_t + \phi_1 INT_{t-1} + \phi_2 INT_{t-2} + \varepsilon_t \\
\log(\sigma^2_t) = \omega + \gamma \log(\sigma^2_{t-1}) + \delta \frac{\varepsilon_{t-1}}{\sigma_{t-1}} + \kappa \frac{\varepsilon_{t-2}}{\sigma_{t-2}}
\]
The coefficient on contemporaneous intervention is always significant and incorrectly signed. The estimates are around -0.016 which in absolute value is much larger than the coefficients from the lag intervention. The wrong sign is a direct consequence of the endogeneity. The OLS estimate is a combination of $\alpha$ and $\beta$, and likely $\gamma$, and can be negative even for positive $\alpha$ if the RBA leans against the wind ($\beta < 0$), or common shocks affect the two equations with different signs ($\gamma < 0$).\textsuperscript{16}

2.5.1 Contemporaneous effects of Central Bank Intervention

We now move on to the results from our basic model (equation (2.10)) that excludes lags. This implies that the central bank target is the previous day’s exchange rate, although the rich error structure will capture deviations from this. Later we extend the model to include lags that implies a target of a longer moving average.

In Table 2.4 we report the coefficients of interest: $\alpha$, $\beta$, $\gamma$, the two thresholds, the constants and the variances. The first column is the point estimate using the simulated GMM methodology and the moments from the data. The second column is the mean of the bootstrapped distribution of the parameter. As mentioned before, we generate 100 sets of moments bootstrapping the residuals of the original data in each of the regimes. The standard deviation is of the bootstrapped distribution of parameter estimates. The fourth column is the quasi $t$-statistic calculated as the mean divided by the standard deviation. It is important to highlight that the distributions are not normal, so this statistic should be considered informative, but not conclusive, in terms of the significance of the coefficients. The fifth and sixth column are the maximum and minimum values of the bootstrapped estimates. The last column is the percentage of the observations that are below zero. This is the statistic that we use to determine the significance of the coefficients. We prefer to look at the mass above and below zero rather than the quasi $t$-statistic given that most of the estimates are not normally distributed (this is a p-value).

\textsuperscript{16}If there are no lags in the system the OLS estimate when the exchange rate return is regressed on intervention will be \[ \frac{(\alpha+\gamma)(\alpha+\beta+\gamma)\sigma_i^2+\sigma_r^2+\sigma^2}{(\alpha+\beta+\gamma)^2}\]
Table 2.4: Estimates of the Standard Model.

<table>
<thead>
<tr>
<th>Exchange rate equation</th>
<th>Point Estimate</th>
<th>Mean Distribution</th>
<th>Standard Deviation</th>
<th>Mean/StDev</th>
<th>Maximum</th>
<th>Minimum</th>
<th>Percentage of observations below zero</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \alpha )</td>
<td>0.1350</td>
<td>0.1364</td>
<td>0.0372</td>
<td>3.66</td>
<td>0.2313</td>
<td>0.0332</td>
<td>0.0%</td>
</tr>
<tr>
<td>( \gamma )</td>
<td>-0.3042</td>
<td>-0.3054</td>
<td>0.0561</td>
<td>-5.44</td>
<td>-0.1457</td>
<td>-0.4481</td>
<td>100.0%</td>
</tr>
<tr>
<td>( c_e )</td>
<td>0.0043</td>
<td>0.0078</td>
<td>0.0188</td>
<td>0.41</td>
<td>0.0904</td>
<td>-0.0859</td>
<td>9.9%</td>
</tr>
</tbody>
</table>

| Reaction function      |                |                   |                    |            |         |         |                                     |
| \( \beta \)            | -0.0279        | -0.0894           | 0.0952             | -0.94      | -0.0047 | -0.5774 | 100.0%                              |
| \( c_{INT} \)          | -0.0909        | -0.1104           | 0.0552             | -2.00      | 0.0232  | -0.358  | 99.0%                               |

| Other parameters       |                |                   |                    |            |         |         |                                     |
| \( INT_t^{1} \)        | 0.3160         | 0.3511            | 0.0869             | 4.04       | 0.9906  | 0.2112  | 0.0%                                |
| \( INT_{k}^{2} \)      | 1.8529         | 1.9767            | 1.9874             | 0.99       | 21.6319 | 1.2606  | 0.0%                                |
| \( \sigma_{x}^{2} \)   | 0.0345         | 0.0159            | 0.0142             | 1.12       | 0.0383  | 0.0000  | 0.0%                                |
| \( \sigma_{y}^{2} \)   | 0.4395         | 0.5025            | 0.1837             | 2.73       | 1.3248  | 0.2381  | 0.0%                                |
| \( \sigma_{e}^{2} \)   | 0.4094         | 0.4562            | 0.1825             | 2.50       | 1.2857  | 0.0652  | 0.0%                                |

Our main focus is on the estimate of \( \alpha \). The point estimate is 0.135, and the mean of the distribution is 0.136. The standard deviation is 0.037. Indeed, as can be seen, from all 100 realisations the minimum is 0.033, which implies that there is no mass below zero. Both the quasi t-statistic and the proportion of realisations below zero indicate that the estimate is highly significant. The distribution of \( \alpha \) is shown in Figure 2-2.

This estimate of \( \alpha \) indicates that intervention has a large effect on the exchange rate. The coefficient on contemporaneous intervention, \( \alpha \) (0.136), implies that $A100m of purchases of Australian dollars is associated (on average) with a 1.36 per cent appreciation of the Australian dollar. If we take the average exchange rate over our sample period to be 0.75, then $US100m of purchases of Australian dollars would have appreciated the Australian dollar by approximately 1.81 per cent. This response is larger than most results in the literature but closely related to calibrations obtained by Dominguez and Frankel (1993c). Dominguez and Frankel calculated that $US100m of purchases of US dollars would appreciate the US dollar by just under 1.6 per cent.

Our estimate differs from that in Dominguez and Frankel (1993c) in that it is calculated directly from exchange rate and intervention data. In constructing their estimate Dominguez and Frankel need to make several assumptions, such as mean-variance preferences of investors, and use expectations survey data and assets supplies data that are likely to contain measurement errors. The dependence on survey data also requires that they consider exchange rate returns over a longer horizon and so can’t estimate the short-run impact of intervention. A significant
improvement of our estimation methodology is that we are able to include both contemporaneous intervention and exchange rate returns to address the simultaneity. Nevertheless, our estimate is close to the one found by Dominguez and Frankel (1993c), and we cannot reject the hypothesis that they are the same.

The point estimate of $\beta$ is -0.028, with mean of -0.089 and a relatively large standard deviation of 0.095. As can be seen in Table 2.4 the quasi t-statistics suggest that the estimate is not statistically significant from zero. However, demonstrating the importance of the non-normality of the distribution, all realisations are below zero indicating the estimate is highly significant. Figure 2-3 depicts the distribution of $\beta$. As can be seen, the distribution is not normal and there are some realisations that are relatively large (close to zero) but the distribution has a long left tail.

The negative coefficient conforms with our priors, and RBA statements, that interventions are dictated by leaning against the wind. The point estimate implies that a one percent unexpected depreciation of the exchange rate leads the central bank to lean against the wind,
buying 28 million Australian dollars in the foreign exchange market to slow the depreciation.

Our estimate of \( \gamma \) is precisely estimated; it is negative and significant. The point estimate is \(-0.304\), almost identical to the mean of the bootstrapped distribution, \(-0.305\), while the standard deviation is 0.056. In this case 100 percent of the bootstrapped estimates are negative, implying that the coefficient is significantly different from zero at high levels of significance. The negative sign of \( \gamma \) demonstrates that shocks common to the exchange rate process and the central bank’s reaction function simultaneously weaken (strengthen) the currency and increase (decrease) the central bank’s desire to support the currency. Again this supports our priors of central bank policy. These shocks can be interpreted to be shocks to the exchange rate during periods of high conditional volatility which the central bank shows greater inclination to resist. The negative estimate of \( \gamma \) provides an additional reason why failing to account for the simultaneity of exchange rate and intervention determination is likely to bias downward the impact of intervention on the exchange rate.

Finally, the estimates of the thresholds are highly significant and reasonable given what we have observed. These values have been scaled to simplify the estimation process. Their
interpretation is that in the early period the central bank participates in the foreign exchange rate market when the shadow intervention is larger than $A3.2m, while in the later regime it intervenes if the required intervention is larger than $A18.5m.

### 2.5.2 Dynamic impact of Central Bank Intervention.

We now extend the model to allow for two lags in the structural equations. Doing so provides a richer representation of the policy function and allows us to examine the temporal impact of interventions on the exchange rate.

The expanded model is

\[
\Delta e_t = c_e + \alpha INT_t + \\
\lambda_{ee,1} \Delta e_{t-1} + \lambda_{eINT,1} INT_{t-1} + \lambda_{ee,2} \Delta e_{t-2} + \lambda_{eINT,2} INT_{t-2} + \\
\gamma z_t + \varepsilon_t
\]

\[
INT^*_t = c_{INT} + \beta \Delta e_t + \\
\lambda_{INT,e,1} \Delta e_{t-1} + \lambda_{INTINT,1} INT^*_t + \lambda_{INT,e,2} \Delta e_{t-2} + \lambda_{INTINT,2} INT^*_t + \\
z_t + \eta_t
\]

where the additional coefficients are the $\lambda$'s. There are 14 coefficients to be estimated. We expand the moments used in the estimation to include across time statistics: we include the probability of two and three consecutive interventions, and the first and second serial correlation in the exchange rate for both the high and low thresholds.\(^{17}\) The results from the estimation are shown in Table 2.5.

\[^{17}\text{We have run the simulation also including other moments such as the probability of two positive interventions, and the correlation in the quantities of two consecutive interventions. The results were not sensitive to these changes.}\]
Table 2.5: Estimates of the Model with Lags.

<table>
<thead>
<tr>
<th>Exchange rate equation</th>
<th>Point Estimate</th>
<th>Mean Distribution</th>
<th>Standard Deviation</th>
<th>Mean/StDev</th>
<th>Maximum</th>
<th>Minimum</th>
<th>Percentage of observations below zero</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\alpha$</td>
<td>0.0446</td>
<td>0.0533</td>
<td>0.0375</td>
<td>1.42</td>
<td>0.1579</td>
<td>0.0011</td>
<td>0.0%</td>
</tr>
<tr>
<td>$\gamma$</td>
<td>-0.0612</td>
<td>-0.1400</td>
<td>0.0844</td>
<td>-1.66</td>
<td>0.0021</td>
<td>-0.4130</td>
<td>99.0%</td>
</tr>
<tr>
<td>$\lambda_{x1}$</td>
<td>-0.0638</td>
<td>-0.0603</td>
<td>0.0348</td>
<td>-1.73</td>
<td>-0.0263</td>
<td>-0.3196</td>
<td>100.0%</td>
</tr>
<tr>
<td>$\lambda_{x2}$</td>
<td>-0.0379</td>
<td>-0.0385</td>
<td>0.0326</td>
<td>-1.18</td>
<td>0.0085</td>
<td>-0.2962</td>
<td>99.0%</td>
</tr>
<tr>
<td>$\lambda_{x INT1}$</td>
<td>0.0369</td>
<td>-0.0007</td>
<td>0.0351</td>
<td>-0.02</td>
<td>0.0866</td>
<td>-0.0856</td>
<td>65.0%</td>
</tr>
<tr>
<td>$\lambda_{x INT2}$</td>
<td>0.0246</td>
<td>0.0021</td>
<td>0.0321</td>
<td>0.07</td>
<td>0.1060</td>
<td>-0.0844</td>
<td>57.0%</td>
</tr>
<tr>
<td>$c_x$</td>
<td>0.0004</td>
<td>0.0100</td>
<td>0.0475</td>
<td>0.21</td>
<td>0.2608</td>
<td>-0.1263</td>
<td>32.0%</td>
</tr>
</tbody>
</table>

| Reaction function      |                |                   |                    |            |         |         |                                       |
| $\beta$               | -0.0657        | -0.1538           | 0.1142             | -1.35      | 0.0000  | -0.6292 | 100.0%                                |
| $\lambda_{INTx1}$     | 0.0140         | -0.0184           | 0.0186             | -0.99      | 0.0604  | -0.0742 | 87.0%                                 |
| $\lambda_{INTx2}$     | -0.0030        | -0.0189           | 0.0149             | -1.27      | 0.0760  | -0.0558 | 94.0%                                 |
| $\lambda_{INT INT1}$  | -0.2299        | -0.0173           | 0.1499             | -0.12      | 0.2819  | -0.6120 | 43.0%                                 |
| $\lambda_{INT INT2}$  | -0.0923        | -0.0155           | 0.1064             | -0.15      | 0.1913  | -0.3556 | 48.0%                                 |
| $c_{INT}$             | -0.0854        | -0.1933           | 0.1150             | -1.68      | 0.0636  | -0.6258 | 98.0%                                 |

| Other parameters       |                |                   |                    |            |         |         |                                       |
| $INT_1$               | 0.3253         | 0.4194            | 0.1007             | 4.16       | 1.0285  | 0.2500  | 0.0%                                  |
| $INT_2$               | 2.6664         | 1.6935            | 0.4353             | 3.89       | 2.7410  | 0.7022  | 0.0%                                  |
| $\sigma_{\epsilon}$  | 0.0128         | 0.0062            | 0.0046             | 1.36       | 0.0200  | 0.0000  | 0.0%                                  |
| $\sigma_{\zeta}$     | 0.3423         | 0.6430            | 0.1495             | 4.30       | 0.9579  | 0.2443  | 0.0%                                  |
| $\sigma_{\eta}$      | 0.3044         | 0.4392            | 0.1283             | 3.42       | 0.7659  | 0.0895  | 0.0%                                  |

The estimates of $\alpha$, $\beta$ have the same signs as those in the base model. The coefficient $\alpha$ is smaller but the coefficients on lagged interventions, $\lambda_{x INT1}$ and $\lambda_{x INT2}$, are also positive, indicating that intervention continues to affect the exchange rate on the days subsequent to the intervention. Notably, these subsequent effects are smaller. The largest impact from an intervention occurs on the day it is conducted. The coefficient $\beta$ is again negative, though larger in magnitude. The importance of accounting for the endogenous relationship is highlighted by the much smaller role that exchange rate changes from previous days have in determining intervention, as seen by the small coefficients on lagged exchange rate returns. The coefficients on lagged intervention are negative due to the tendency of subsequent interventions to be smaller than the initial intervention. The bootstrapped distributions are depicted in Figures 2-4 and 2-5.

The interpretation of the response of the exchange rate to intervention is complicated by the feedback between the variables in the presence of lags. Figure 2-6 shows the impulse response functions for the exchange rate in the model with lags and the base model without lags. The impulse responses for the two models are calculated ignoring the participation decision of the central bank. Since the coefficients in the policy function on lagged intervention are negative,
Figure 2-4: Bootstrapped distribution of $\alpha$ when lags are included in the model.

Figure 2-5: Bootstrapped distribution of $\beta$ when lags are included in the model.
abstracting from the participation decision implies the central bank will unwind its intervention on subsequent days. Empirically this does not occur demonstrating the importance of the interaction of the two components of the central bank policy function. The central bank almost never intervenes on opposite sides of the market on subsequent days. A third series on the graph incorporates the entry decision by restricting the central bank to remain absent from the market on the days following the initial intervention. Here the effect on subsequent days is shown to be larger, though still smaller than that on the initial day.

The cumulative impact of intervention is shown in Figure 2-7. Note that the short lag structure used constrains that intervention has a permanent effect on the exchange rate. A long lag structure would be needed to capture the unwinding of the effects of intervention. Including many lags is not practical for the estimation method we employ and so we are unable to assess how long the effects of intervention last beyond a few days. The cumulative response of the exchange rate to intervention, after 2 to 3 days, is similar for the models including and excluding lags. In the model with lags $A100m ($US100m) will appreciate the exchange rate
Figure 2-7: Impulse response: increase in level of exchange rate from $A100 million intervention shock.

by 1 per cent (1.35 per cent), slightly less than the 1.36 per cent (1.81 per cent) predicted by the model excluding lags.

2.5.3 Caveats and further research

Before concluding, it is worth spending time discussing the robustness and validity of the procedure developed in this paper. Our estimation methodology has several caveats that should be addressed in future research. First, it does not distinguish through which channel intervention affects the exchange rate. Given the relative shifts in bonds in private hands is small, the portfolio balance channel is likely to play an insignificant role. Rather, the signalling channel, and possibly market microstructure effects, would seem to be responsible - but this is a conjecture, rather than a result.

Second, as it stands our methodology requires that the change in policy is truly exogenous. Finding such exogenous changes in policy for other countries may not be straight forward. However, it is possible to model the threshold in the participation equation as a function of
macro variables, the exchange rate, and even the second moments of the endogenous variables. The procedure of estimation would be exactly the same as the one described here if the change in policy is thought as a shift in the coefficients in that equation. Furthermore, the participation decision could be rationalised as a switching Markov regime where the transition probabilities are function of the endogenous variables. The estimation of these models is beyond the scope of the present paper, and are left for future research. Nevertheless, it is important to mention that in this paper we have emphasised the identification issue, and our specification is a reduced form representation of these models where the shift in policy would be reflected in changes in the intervention decision equation.

Finally, our methodology assumes that the coefficients are stable across regimes (other than the threshold). This seems to be a reasonable assumption in the current study but may be more difficult to justify in other contexts. This criticism is akin to the application of the Lucas critique to models of monetary policy. Our framework is not entirely exempt from this critique, and therefore, conclusions from this analysis are subject to the caveat of the strength of these assumptions.

2.6 Conclusion

The endogeneity of exchange rates and intervention has long plagued studies of the effectiveness of central banks' actions in foreign exchange markets. Researchers have either excluded contemporaneous intervention, so that their explanators are predetermined, or obtained a small, and typically incorrectly signed, coefficient on contemporaneous intervention. Failing to account for the endogeneity, when central banks lean against the wind and trade strategically, will likely result in a large downward bias to the coefficient on contemporaneous intervention – explaining the negative coefficient frequently obtained.

This paper uses a novel identification assumption, a change in RBA intervention policy, that allows us to estimate a model that includes the contemporaneous impact of intervention. We use simulated GMM to estimate the model. There are three main results. Our point estimates suggest that central bank intervention has a substantial effect. We find that a (sterilised) purchase of $US100m of Australian dollars by the RBA would be associated with an appreciation
of between 1.35 and 1.81 per cent. These estimates are remarkably similar to that in Dominguez and Frankel (1993c) even though our estimation methodology is completely different, depending on only exchange rate and intervention data. Second, an intervention is shown to have its largest effect on the exchange rate on the day in which it is conducted, with smaller effects on subsequent days. This finding has not previously been demonstrated in the literature due to the problem of endogeneity, and confirms the beliefs of central banks of the immediacy of the efficacy of intervention. Finally, we confirm findings that Australian central bank intervention policy can be characterised by leaning against the wind.
Bibliography


Chapter 3

Modeling commodity currencies:
'The Aussie and two birds'

3.1 Introduction

This paper estimates a dependent economy model of the exchange rate for the Australian, Canadian and New Zealand dollars, the 'Aussie', the 'loony' and the 'kiwi'. The dependent economy model applies to small open economies who are price takers in international trade and, as such, is well suited to these small commodity exporting countries. In this model the exchange rate is determined by balance of payments equilibrium, in the spirit of Nurske (1945), so that the fundamentals of the exchange rate are trade prices, as well as domestic money supply and non-traded output. The essence of this framework is sometimes applied to real exchange rates but is shown here to be useful for nominal exchange rates.¹ The estimated model provides a good representation of the exchange rate for all three countries. The models for the Australian and New Zealand dollars use the commodity price index to proxy for the export price and so may be described as commodity currencies. The Canadian dollar model using a commodity price index is not well specified. Rather, the full export price index is needed to model the

¹Neary (1988) describes the determinants of the real exchange rate for balance of payments equilibrium in a multi good setting while Chen and Rogoff (2001), Clinton (2001), Djouciad et al. (2000) and Gruen and Wilkinson (1994) estimate models of the real exchange rate inspired by this framework. Sjaastad (1998) applies a similar methodology to both the real and nominal Swiss franc exchange rates. Freebairn (1990) and Sjaastad (1990) also consider whether the Australian dollar is a commodity currency.
exchange rate. As such the Canadian dollar’s place as a commodity currency is more tenuous.

The models are used to address the finding of Meese and Rogoff (1983) that the out-of-sample projections of exchange rate models are less accurate than a random walk inspired no-change projection. As documented by Frankel and Rose (1995) this result has largely endured the attempts of many researchers to overturn it.²

For the two larger economies the model of the exchange rate is found to outperform a random walk at most horizons. As with the existing research on exchange rates this test uses the realised values of explanatory variables to focus attention on the estimated model rather than the forecasts of the explanators. The New Zealand dollar model breaks down early in the out-of-sample period, at about the time of the Asian crisis, suggesting this framework is not completely robust to financial factors.

The paper then goes beyond the Meese and Rogoff finding by considering forecasts based solely on past and current information. An appropriately weighted commodity price index is found to be an excellent proxy for the price of exports for Australia. Commodity futures can then be used to measure expectations of future values of one of the main exchange rate fundamentals. A simple model that uses only current and past information, including commodity futures prices, to forecast the exchange rate is found to be more accurate than an assumption of no-change in the exchange rate. Exchange rates are notoriously volatile and the forecasts generated do have large errors. But the finding that exchange rates are in part predictable using economic fundamentals, at least in this special case where fundamentals can easily be measured, is promising.

The rest of the paper is as follows. Section 3.2 briefly describes the dependent economy model. Section 3.3 then discusses the data available to estimate the model and the nature of the three countries used in the study. The estimation and results, including the out-of-sample projections and forecasts, are presented in Section 3.4. Section 3.5 concludes and discusses implications of the findings.

²Almost all of this empirical research has been based upon standard monetary models of the exchange rate. For a summary of these models, and copious references, see Taylor (1995). The performance of exchange rate models is considered in several papers in the forthcoming February 2003 issue of the Journal of International Economics to commemorate 20 years since the Meese and Rogoff paper.
3.2 The dependent economy model of the exchange rate

Chapter 1 outlined a simple dependent economy model of the exchange rate. The dependent economy is highly integrated with other countries and is a price taker in world markets. It exports a good of which it has a fixed per period endowment, a reasonable assumption for commodity exporting countries, and imports a traded consumption good. Domestic agents also eat a non-traded consumption good, of which the economy is endowed with a random quantity each period. The nominal exchange rate also depends on the domestic money stock. If expectations of the future money stock are equal to today's value, the exchange rate equation simplifies to

\[ s_t = m_t - \left( \gamma + \frac{\rho}{\delta} (1 - \gamma) \right) p^*_X + (1 - \gamma) \left( \frac{\rho}{\delta} - 1 \right) p^*_M - \gamma \left( \frac{\rho}{\delta} + 1 \right) y_N + \text{const} \]  

(3.1)

where \((1 - \gamma)\) is the degree of openness of the economy (the weight on the traded consumption good in the Cobb-Douglas aggregate good), and \(\rho\) and \(\delta\) are the intertemporal elasticity of substitution and the elasticity of money demand. Clearly the exchange rate will be a negative function of the foreign currency price of the export good, \(p^*_X\). The sign of the coefficient on the foreign currency price of the traded, or imported, good, \(p^*_M\), will depend on the parameter values.\(^3\) The exchange rate should also decline in non-traded output, \(y_N\), in order for the price of the traded good to adjust to restore internal balance. In the simple representation given in equation (3.1) in which money shocks are fully persistent the exchange rate moves one-for-one with the contemporaneous domestic money stock. In the model the money stock is assumed to be set exogenously by the monetary authority, but this decision could potentially depend on the exogenous variables; the foreign prices and non-traded output. The monetary authority could also consider the exchange rate in setting the money stock, although in practice the countries examined here have freely floating exchange rates so that the central bank does not set policy with respect to contemporaneous exchange rate changes.\(^4\) As such the money stock should be predetermined with respect to the exchange rate. Further, in the estimation a broad money

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\(^3\) The ambiguity of the sign is because of two offsetting effects through real tradable wealth and real money demand.

\(^4\) The exception is the NZ experiment with setting monetary policy based on a monetary conditions index, as discussed below.
aggregate is used which is beyond the direct control of the central bank.

The model makes several simplifying assumptions. Counterfactually, the law of one price is assumed to hold in order to define the exchange rate. The assumption about random walk prices implies that the country will not use international markets to smooth consumption and so trade is always balanced. Further, the potential for output response to shocks through endogenous labour supply or real effects from monetary shocks through sticky prices are ignored. Many of these factors are likely to be much less relevant over longer periods and so equation (3.1) should provide a good representation of the long-run equilibrium exchange rate. The empirical estimation will attempt to account for these factors by modeling the short-run dynamics of the system.

3.3 Data

The model is estimated using monthly data which requires the use of some proxy series. All prices are measured in US dollars and exchange rates are bilateral rates of units of the national currency per US dollar. The data Appendix describes the data in more detail and provides sources for all series. The samples used are based on the period of floating exchange rates, from December 1983 for Australia, May 1970 for Canada and March 1985 for New Zealand. The Canadian sample is shortened to start in January 1975 to account for an initial period of uncertainty regarding the behaviour of floating exchange rates. The New Zealand sample is constrained to start in January 1986 by the availability of data.

<table>
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<th>1980s</th>
<th>1990s</th>
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</thead>
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<td>75</td>
<td>73</td>
<td>60</td>
</tr>
<tr>
<td>Canada</td>
<td>60</td>
<td>51</td>
<td>40</td>
</tr>
<tr>
<td>New Zealand</td>
<td>65</td>
<td>61</td>
<td>52</td>
</tr>
</tbody>
</table>

Australian and New Zealand export price data are only available at a quarterly frequency. For both of these countries well over half of their exports are commodities, as seen in Table 3.1, although this has declined. Individual commodities are largely homogenous and typically traded on open markets, and so their prices are available at higher frequencies. An index of
commodity prices can then serve as a proxy for the price of exports. The Reserve Bank of Australia produces a monthly commodity price index weighted by export shares. As seen in panel 1 of Figure 3-1 this index is an excellent proxy for the quarterly export price series. Figure 3-3 shows that ANZ bank’s New Zealand export-weighted commodity price index also has a close relationship with export prices, though not as tight as that for Australia. The weaker relationship may result from the New Zealand series being constructed from non-official sources and the larger proportion of agricultural, and so less homogenous, exports compared with Australia. While the proportion of commodities in Canada’s exports is smaller than that of the antipodean countries, as Figure 3-2 shows, it is still clearly a very important component of changes in the price of exports, which are available monthly. The dependent economy model is still likely to be valid for Canada since it is very open and over three-quarters of its trade is with the United States, an economy over eleven times larger.5

The second panel of Figures 3-1, 3-2 and 3-3 demonstrates the relationship between the exchange rate and commodity prices over the period of floating exchange rates. Figure 3-

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5These are averages over the sample period 1975-2001. The proportion of trade with the US has increased steadily and is now over 85 per cent.
Figure 3-2: Canada: commodity prices, export prices and the exchange rate.

Figure 3-3: New Zealand: commodity prices, export prices and the exchange rate.
I demonstrates financial market wisdom that the Australian dollar moves very closely with the Australian commodity price index. The New Zealand dollar has a similar relationship, though both show a large depreciation at the end of the sample. The longer Canadian sample demonstrates more vividly the depreciating trend in the exchange rate, that also exists for Australia and New Zealand. The dependent economy model would suggest that this is the result of higher domestic inflation and possibly a decline in the relative price of export goods.\(^6\)

As with export prices, import prices are available on a monthly basis for Canada but not for Australia or New Zealand. Since both countries are largely self sufficient in non-oil commodities their imports consist mostly of manufactured goods. As a proxy for their import prices the price of exports from industrial countries is used. Because they are also oil dependent the price of oil is included along with the that of exports from industrial countries.

All three economies experienced a large degree of financial reform and deregulation over the course of the sample used, resulting in rapid growth in the monetary aggregates. Narrower aggregates were more affected as a higher proportion of funds was progressively captured by these measures. To mitigate the influence of these shocks, the broadest monetary measures available are used. For Australia this is broad money, for Canada it is M2++ and for New Zealand it is M3 broad money. The dependent economy model is stationary in that it does not account for expansion of the economy, which leads to a larger money stock for given price levels. To account for this, the money stock is divided by real GDP to keep it on a unit basis consistent with the model.

The unemployment rate is used to proxy non-traded output since much of the non-traded sector consists of the labour intensive services sector and it is available on a monthly basis for Australia and Canada. The New Zealand series is interpolated from a quarterly frequency.

3.3.1 Stationarity tests

The variables in the dependent economy model are typically non-stationary requiring appropriate econometric techniques be employed. The Elliot, Rothenberg and Stock (ERS) (1996)

\(^6\)This would be consistent with the Prebisch-Singer hypothesis that there is a long-term downward trend in primary commodity prices relative to manufactured goods, due to the low income elasticity of demand for commodities and rapid increases in supply, but empirical evidence on this is very sensitive to the period considered.
test for a unit root, which quasi-differences the data before employing an ADF test, is used as this has greater power than standard unit root tests. The lag length is selected based on the BIC as Phillips and Xiao (1998) demonstrate this increases the test’s power. All variables apart from the unemployment rate are in logs. The test statistics in Table 3.2 shows that the presence of a unit root cannot be rejected in the levels of any of the variables, with the possible exception of oil, in the Australian sample. A tilde over the import price is used to indicate that this is a proxy series, \( \hat{p}_{M} \), while \( p_{Xc}^{*} \) is used to indicate the commodity export price series. The existence of a unit root in the first differences is rejected for all of the nominal variables at the 1 per cent level of significance, and for the unemployment rate at the 5 per cent significance level. The results for New Zealand are similar, though the rejection of a unit root in the first differences is weaker, and again the price of oil is potentially stationary. The test results on the Canadian variables are less resounding with the hypothesis of a unit root in the first differences of money and the import price, \( p_{M}^{*} \), and export price, \( p_{X}^{*} \), not rejected at conventional significance levels. This is despite the rejection of a unit root in the first difference of the commodity price index, \( p_{Xc}^{*} \), and the commodity price index excluding oil, \( p_{Xcov}^{*} \).

Unit root tests have low power and so the failure to reject the existence of a unit root is not necessarily conclusive evidence that the variable is I(1). As a guard against this low power the Kwiatkowski, Phillips, Schmidt and Shin (KPSS) (1992) test, which has a null of stationarity, is employed. These results are shown in Table 3.3. For all three countries the KPSS test, for various truncation lags of the Bartlett window, provides strong evidence against trend stationarity of the levels of the variables. With few exceptions the stationarity of the first difference of the variables is not rejected. For Australia and Canada the stationarity of the first difference of money is rejected. This seemingly results from the effective break in the series with the sharp reduction in inflation in the early 1990s. Applied to subsamples before and after the introduction of inflation targeting there is much greater evidence of first-difference stationarity. So long as this change does not affect the structural parameters of the model, the model should be able to account for this break in monetary growth through changes in the dependent variable. That the exchange rate is found to be I(1) despite this possible break highlights the fragility of unit root tests in small samples. The stationarity of the first difference of the Canadian import price is rejected, though if a trend is included it is not rejected even at high confidence.
Table 3.2: Elliot, Rothenberg and Stock (1996) test for unit root.

<table>
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</thead>
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<tr>
<td></td>
<td>including trend</td>
<td>no trend</td>
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<tr>
<td></td>
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</tr>
<tr>
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<tr>
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# Lag length selected by BIC

***, **, * indicate significance at the 1,5,10% level

Critical values from Elliot et al (1996)

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<tr>
<td>10%</td>
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levels. The need to include a trend possibly results from the longer Canadian sample which includes the decline in inflation from the high levels of the late 1970s. The stationarity of the first difference of the New Zealand unemployment rate is rejected though this is not surprising given the series is interpolated. With no alternative monthly indicator of non-traded output this series is retained.

3.4 Estimation

As seen in Section 3.3.1 the variables suggested by the dependent economy model to be explainators of the exchange rate are non-stationary. If they do explain the behaviour of the exchange rate, at least over the long run, these variables should be cointegrated with the exchange rate. The Phillips and Hansen (1990) fully-modified OLS (FM-OLS) estimator is used to efficiently estimate this long-run relationship involving I(1) series. Hansen (1992) modifies this estimator to account for drift in the exogenous variables and a trend in the cointegrating vector. Estimation of the long-run relationship is conducted using the Coint package for Gauss. The use of a single equation framework is justified by the exogeneity of the foreign price series that is a crucial part of the dependent economy framework. Non-traded output will also be at least weakly exogenous, and given the broad monetary aggregates used and the delayed response of central bank policy to the exchange rate, the money stock should also be exogenous to the exchange rate.

The short-run dynamics are modeled with the deviations from the long-run equilibrium in a parsimoniously parameterised error-correction framework. This parametric approach to short-run changes will be useful to account for the factors mentioned earlier that may cause temporary deviations from the long-run equilibrium relationship.\(^7\)

The models are initially estimated up to December 1995 to allow testing of their out-of-sample properties. The results from the FM-OLS estimation of the long-run relationship for Australia and New Zealand are shown in Table 3.4. For both countries the model is estimated including and excluding the price of oil. A trend included to account for the growth in money from financial innovation and reform has a small coefficient and is insignificant and so is not

\(^7\)Phillips and Loretan (1991) advocate this approach of estimating the long-run relationship efficiently and then using a parametric methodology for the short-run dynamics.
Table 3.3: Kwiatkowski, Phillips, Schmidt and Shin (1992) test for stationarity.

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</tr>
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<td>0.48***</td>
<td>0.29***</td>
<td>0.22***</td>
<td>0.18**</td>
</tr>
<tr>
<td>$m$</td>
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<td>1.00***</td>
<td>0.58***</td>
<td>0.42***</td>
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<td>0.42***</td>
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<td>0.24***</td>
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<td>$p_{M}$</td>
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<td>0.12</td>
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<td>0.54**</td>
<td>0.50**</td>
<td>0.42**</td>
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<tr>
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<td>0.13</td>
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<td><strong>Canada</strong></td>
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<tr>
<td>Levels - inc trend</td>
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</tr>
<tr>
<td>$s$</td>
<td>2.68***</td>
<td>0.69***</td>
<td>0.40***</td>
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<tr>
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<td>0.54**</td>
<td>0.72**</td>
<td>0.69**</td>
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<td>0.53**</td>
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</tr>
<tr>
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<td>0.52***</td>
<td>0.31***</td>
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<td>0.18**</td>
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<td>0.21**</td>
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<td>3.42***</td>
<td>0.87***</td>
<td>0.51***</td>
<td>0.36***</td>
<td>0.29**</td>
</tr>
<tr>
<td>$p_{ot}$</td>
<td>0.68***</td>
<td>0.19**</td>
<td>0.13*</td>
<td>0.10</td>
<td>0.09</td>
</tr>
<tr>
<td>First difference - no trend</td>
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<td>$s$</td>
<td>0.37*</td>
<td>0.40*</td>
<td>0.38*</td>
<td>0.35*</td>
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<td>0.15</td>
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<td>0.14</td>
</tr>
<tr>
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<td>0.24</td>
<td>0.25</td>
<td>0.24</td>
</tr>
<tr>
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<td>0.65**</td>
<td>0.53**</td>
<td>0.46*</td>
<td>0.40*</td>
</tr>
<tr>
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<td>0.07</td>
<td>0.08</td>
<td>0.09</td>
<td>0.09</td>
</tr>
</tbody>
</table>

***, **, * indicate significance at the 1,5,10% level

Critical values from KPSS (1992)

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</tr>
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<tr>
<td>5%</td>
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</tr>
<tr>
<td>10%</td>
<td>0.119</td>
</tr>
</tbody>
</table>

103
included. The model performs well in explaining the Australian dollar - US dollar exchange rate. The money stock is significant with the expected positive sign. Though the model would suggest a coefficient of unity the coefficient is significantly less than unity, likely the result of financial reform affecting monetary aggregates and changing velocity of money. The export price (the commodity price) is highly significant and negative as expected. The coefficient on import prices is very small and insignificantly different from zero suggesting the series used may not be a good proxy for the actual price of imports. The negative significant sign on the unemployment rate is puzzling. An increase in the unemployment rate should be associated with a decline in non-traded output and so a depreciation to restore internal balance. Unemployment may be a poor proxy of the non-traded output as employment typically lags output. But with no obvious alternative monthly measure of non-traded output it is retained. The model is robust to the inclusion of the price of oil, which is marginally significant and does little to improve the fit. The Elliot, Rothenberg and Stock test for a unit root used in Section 3.3.1 is applied to the residuals from the long-run relationship as a test of cointegration. The null of a unit root is resoundingly rejected indicating the series are cointegrated.

In the New Zealand model the trend to compensate for financial reform is significant and results in a positive and significant coefficient on money and so it is included. Despite the apparent correlation seen in Figure 3-3 the export price has a relatively small and insignificant coefficient. The import price is significant, though the magnitude of the negative coefficient is surprising. The coefficient on the unemployment rate is however positively signed and significant. Again the inclusion of the price of oil does not improve the model's fit or alter the other coefficients. The rejection of a unit root in the residuals demonstrates that the explanators are cointegrated with the exchange rate.

The bottom half of Table 3.4 contains the results from the estimation of the short-run dynamics in the error-correction model (ECM). The lagged residuals from the long-run relationship, the error-correction term, $\Delta c$, which represent the distance from equilibrium, is included in a regression of the one-month change in the exchange rate on contemporaneous and lagged changes in the independent variables and lagged changes in the exchange rate. Insignificant lags are removed to arrive at a parsimonious model of the short-run relationship. The coefficient on the error-correction term is significantly negative indicating the dynamics return the system in
<table>
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<tr>
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<th>New Zealand</th>
</tr>
</thead>
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<td></td>
</tr>
<tr>
<td></td>
<td>-0.002***</td>
<td>-0.002***</td>
</tr>
<tr>
<td></td>
<td>(0.001)</td>
<td>(0.001)</td>
</tr>
<tr>
<td><strong>$m$</strong></td>
<td>0.465***</td>
<td>0.418***</td>
</tr>
<tr>
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<td>(0.150)</td>
<td>(0.138)</td>
</tr>
<tr>
<td><strong>$p_{Xc}^*$</strong></td>
<td>-1.083***</td>
<td>-0.976***</td>
</tr>
<tr>
<td></td>
<td>(0.132)</td>
<td>(0.125)</td>
</tr>
<tr>
<td><strong>$\tilde{p}^*_{M}$</strong></td>
<td>-0.002</td>
<td>-0.039</td>
</tr>
<tr>
<td></td>
<td>(0.185)</td>
<td>(0.176)</td>
</tr>
<tr>
<td><strong>$unemp$</strong></td>
<td>-0.036***</td>
<td>-0.030***</td>
</tr>
<tr>
<td></td>
<td>(0.007)</td>
<td>(0.007)</td>
</tr>
<tr>
<td><strong>$p_{oil}^*$</strong></td>
<td>-0.067**</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.039)</td>
<td>(0.039)</td>
</tr>
<tr>
<td><strong>ERS test of cointegration</strong></td>
<td><strong>test stat</strong></td>
<td><strong>lags</strong></td>
</tr>
<tr>
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<td>-4.179***</td>
<td>0</td>
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<td></td>
</tr>
<tr>
<td><strong>Short-run dynamics: dependent variable $\Delta s_t$</strong></td>
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<td></td>
</tr>
<tr>
<td><strong>$const$</strong></td>
<td>0.0002</td>
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<tr>
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<td>(0.003)</td>
<td>(0.003)</td>
</tr>
<tr>
<td><strong>$e_{Ct-1}$</strong></td>
<td>-0.145***</td>
<td>-0.193***</td>
</tr>
<tr>
<td></td>
<td>(0.046)</td>
<td>(0.050)</td>
</tr>
<tr>
<td><strong>$\Delta m_t$</strong></td>
<td>-0.705**</td>
<td>0.261**</td>
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<td></td>
<td>(0.337)</td>
<td>(0.131)</td>
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<tr>
<td><strong>$\Delta m_{t-1}$</strong></td>
<td>-0.720**</td>
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<tr>
<td></td>
<td>(0.331)</td>
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<tr>
<td><strong>$\Delta m_{t-3}$</strong></td>
<td>0.896***</td>
<td>0.807**</td>
</tr>
<tr>
<td></td>
<td>(0.334)</td>
<td>(0.319)</td>
</tr>
<tr>
<td><strong>$\Delta p_{Xc,t}$</strong></td>
<td>-0.780***</td>
<td>-0.661***</td>
</tr>
<tr>
<td></td>
<td>(0.129)</td>
<td>(0.125)</td>
</tr>
<tr>
<td><strong>$\Delta p_{Xc,t-4}$</strong></td>
<td>-0.288***</td>
<td>-0.283***</td>
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<tr>
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<td>(0.103)</td>
<td>(0.103)</td>
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<tr>
<td><strong>$\Delta \tilde{p}^*_{M,t}$</strong></td>
<td>-0.246**</td>
<td>-0.392***</td>
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<td>(0.123)</td>
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<td><strong>$\Delta p_{oil,t-2}$</strong></td>
<td>0.084***</td>
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<tr>
<td><strong>$\Delta s_{t-2}$</strong></td>
<td>0.157**</td>
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<tr>
<td><strong>$\Delta s_{t-4}$</strong></td>
<td>-0.146**</td>
<td>-0.164**</td>
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<td>(0.072)</td>
<td>(0.070)</td>
</tr>
<tr>
<td><strong>$R^2$</strong></td>
<td>0.304</td>
<td>0.362</td>
</tr>
<tr>
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<td>0.229</td>
<td>0.227</td>
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</table>

***, ***, indicate significance at the 1,5,10% level
the direction of the long-run equilibrium. In the Australian model two money growth terms are significant though with opposite signs. The negative coefficient on the first, and positive on the second, of these lags indicates that the exchange rate has a delayed response to money shocks. The contemporaneous change in the commodity price is highly significant with a coefficient close to the long-run value indicating that adjustment to commodity price shocks is rapid. The negative coefficient on lagged dependent variables indicates the presence of some weak negative autocorrelation, or overshooting, that is not captured by the other explanators. Overall the fit is very good with an adjusted-$R^2$ over 0.3.

The short-run dynamics for New Zealand are virtually identical for the two specifications of the long-run relationship. The error-correction relationship is significant and negatively signed demonstrating the importance of disequilibrium in the short-run dynamics. The contemporaneous growth in the money stock has a positive coefficient suggesting the exchange rate responds relatively rapidly to money shocks. While the contemporaneous export price is intuitively signed, its impact is unwound by a lag of four months. The contemporaneous change in the proxy import price again has a puzzling large negative coefficient.

The dependent economy model for Canada is estimated with three different versions of the export price. The first is the Bank of Canada commodity price index, the second is the commodity price index excluding oil, and the third is the actual export price series. While the coefficients on the export price and unemployment rate are intuitively signed they are not significant in the models using the two commodity prices for the export price. The coefficient on the money stock, while insignificant, is negative. In these models only the import price is significant. Not surprisingly the existence of a unit root in the residuals is not rejected using the ERS test.

The results are starkly different using the actual export price. The coefficients are almost all significant and intuitively signed, and the ERS test indicates the existence of a strong cointegrating relationship. The coefficients on the export price and the import price and negative and positive respectively as is suggested by the model. Djoudad et al (2000) suggest there is a special role for oil in their model of the real Canadian dollar exchange rate and so it is also included separately, despite its inclusion as a component of both the export and import price indices. The price of oil is significant and has a positive sign consistent with the net oil
Table 3.5: Dependent Economy Model of Canada.

<table>
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<tr>
<th></th>
<th>Commodity price</th>
<th>Commodity price exc. oil</th>
<th>Export price</th>
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<tr>
<td>$const$</td>
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<td>(0.621)</td>
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<td>(0.197)</td>
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<td>$p^*_Xc$</td>
<td>-0.340</td>
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<td></td>
<td>(0.350)</td>
<td>(0.484)</td>
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</tr>
<tr>
<td>$p^*_Xco$</td>
<td>-0.332</td>
<td>-0.336</td>
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</tr>
<tr>
<td>$p^*_X$</td>
<td>-1.405***</td>
<td>-2.301***</td>
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<td>$p^*_M$</td>
<td>0.872*</td>
<td>0.802</td>
<td>0.692***</td>
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<tr>
<td></td>
<td>(0.526)</td>
<td>(0.508)</td>
<td>(0.241)</td>
</tr>
<tr>
<td>unemp</td>
<td>0.031*</td>
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ERS test of cointegration

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Short-run dynamics

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</tr>
</tbody>
</table>

$R^2$           | 0.038     | 0.036| 0.057        | 0.057        | 0.116        | 0.141        |

***, ***, indicate significance at the 1,5,10% level
importer status of Canada. The money stock is again negative if the price of oil is excluded but is positive and significant in the variant including the oil price. In this second version however the unemployment rate is negative, though with a small coefficient. The substantially better model derived when substituting the actual export price series for the commodity export price series indicates that the commodity series is an incomplete representation of the price Canada faces for it's export goods and so as an explanator of the Canadian dollar exchange rate.

The short-run dynamics are highly parsimonious. Not surprisingly given the first stage results the models using the commodity price proxies of the export price provide a poor fit. The models using the actual export price include the contemporaneous change in the export price with a negative coefficient, indicating that, like the Australian model, the exchange rate adjusts rapidly to export price shocks. When included the change in the price of oil, lagged one period, also shows rapid adjustment in the direction anticipated. Despite the strong long-run relationship, and the significant short-run dynamics, the error-correction term is not significant at conventional levels so that disequilibrium from the long-run relationship does not have a strong-effect on the short-run exchange rate returns.

3.4.1 Out-of-sample projections

A consistent finding of exchange rate research has been the poor performance of empirical models, as documented by Frankel and Rose (1995). Central to this has been the general failure of empirical models to overturn the finding of Meese and Rogoff (1983) that out-of-sample projections from empirical models were less accurate than a random walk inspired no-change projection. More recent use of advanced econometric techniques have had limited success, see for example Mark (1995) and MacDonald and Taylor (1994).

This section presents results for out-of-sample monthly projections using the preferred models estimated in Section 3.4, excluding oil for Australia and including oil for Canada. As noted below the New Zealand model breaks down early in the out-of-sample period and so is not used for out-of-sample projections. The projections are presented for horizons out to 24 months for the period January 1996 to June 2001. Two versions of the model are compared to the no-change projections. The first reestimates the model with the addition of each month's data. The alternative version of the model retains the coefficient values estimated in Section 3.4
using the data up to December 1995. As is conventional in the literature on estimation of exchange rate models the realised values of explanators are used to generate the projections. As such the projections are not forecasts, rather they provide an assessment of the out-of-sample performance of the model.

The versions of the models with updated coefficients were reestimated iteratively with the addition of each extra month of data up to May 2001, a total of 64 times. Since these results are very similar to those presented above they are not shown here. The models for Australia and Canada continue to demonstrate a highly cointegrated long-run relationship over all sample periods. The short-run dynamics are also reestimated each period, with the selection of variables included potentially changing. The only substantive change in the short-run dynamics of the Australian model is the omission of the c \*emporaneous change in the money stock after mid 1998.

The Canadian model is very stable with virtual no change in the estimated parameters from either the short-run or long-run relationships. The significance of the error-correction term increases with the addition of extra data so that for samples ending in a three year window from 1996-99 this term is significant at the 5 per cent level. For models estimated after this the error-correction term is significant at approximately the 10 per cent level or just above.

The New Zealand model breaks down early in the out-of-sample period. The evidence of cointegration in the long-run relationship progressively weakens after 1998 coinciding with the Asian crisis and the large depreciation of the New Zealand dollar. Several of the estimated long-run parameters become insignificant and even change sign at the same time. Since the New Zealand model breaks down early in the out-of-sample period it is not included in the out-of-sample projections.

Figure 3-4 plots the root-mean squared error (RMSE), and mean absolute error (MAE), of projections for Australia. The RMSE penalises large deviations relatively more heavily than does the MAE. The RMSE and MAE for the no-change projections are slightly convex in the horizon.\footnote{If the exchange rate was indeed a random walk so that the k-period variance was k times the one period variance the RMSE and MAE of a no-change projection would increase at rate $\sqrt{k}$ and so be slightly convex in the horizon. This is close to the rate of increase observed in the RMSE and MAE for the no-change projections.} Both the RMSE and MAE show that the no-change projection is marginally more accurate than the model based projections for horizons up to around 8 months. Thereafter
the RMSE and MAE taper off for the model projections and are less than those for the no-change projections. As would be expected the projections are slightly less accurate from the version of the model for which the parameter estimates are not updated. By a horizon of 24 months the errors from projections using the reestimated model are approximately half those of the no-change projections. Obviously the very long horizon, 24 month, projections use fewer independent observations and so these comparisons must be treated with some caution.

The out-of-sample projections for Canada are even more accurate than those for Australia, as seen in Figure 3-5. At all horizons, down to even one month, the projections from both the reestimated and constant parameter versions of the model are more accurate than those based on a no-change assumption. The accuracy of the model projections relative to the no-change projections increases with the horizon of comparison. Again by a horizon of 24 months the errors from the model projections are of the order of half of those based on the no-change assumption. Consistent with the estimated parameters for Canada being slightly more stable than those for Australia the forecast accuracy is less dependent on whether the model is reestimated.

3.4.2 Forecasts

The evaluation of the models’ out-of-sample projections in Section 3.4.1 used the realised values of the dependent variables and so is not a test of the models’ ability to predict exchange rate
movements. In the case of Australia, the commodity price index was found to be an excellent proxy for the export price and as such explained a large portion of changes in the Australian dollar. Since futures and forward contracts exist for many commodities, and their prices are reasonably good predictors of future spot prices, as shown in Chapter 1, they can be used to measure the expected change in one of the major determinants of the Australian dollar. This section exploits this relationship to examine the predictability of the Australian dollar using the commodity price index described in Chapter 1. Because this index is only available for 3-month futures prices the model is reestimated using quarterly data. Data restrictions also limit the total sample to March 1984 to December 2000. Since there are substantially fewer observations, and because of the desire to limit the model to using known variables for forecasting, a highly parsimonious specification of the short-run dynamics inspired by the monthly model is used. Only the contemporaneous change in the commodity price and the lagged error-correction term are included. Data for many of the fundamental explanators are not available at the end of the period they measure and so the model is also estimated lagging the error-correction term two periods.\textsuperscript{9} This is a very conservative timing as many of the monthly series are available with a delay of only one month. The results for the quarterly models estimated with the spot futures

\textsuperscript{9}To generate the error correction term in the forecast period the values of the fundamentals other than commodity prices are held constant at their values from the initial forecast.
commodity price index are shown in Table 3.6.\textsuperscript{10} Since the actual import price series is available on a quarterly basis the model is also estimated using this rather than the proxy import price series. The coefficients in the long-run specification are almost identical to those from the monthly model when using the proxy import price. When the actual import price is used the coefficient on this series is significant and negative, and the coefficient on the commodity price declines in magnitude though it remains highly significant. For both variants, the presence of a unit root in the residuals is rejected by the ERS test at high levels of significance, indicating the fundamentals are cointegrated. The estimated coefficient on the contemporaneous change in the commodity price is not significantly different between the specifications, and is large reaffirming the rapid response of the exchange rate to commodity price shocks. In both cases the error-correction term is highly significant in the ECM, though the point estimate is lower when the error-correction term is lagged two periods.

For comparison with the monthly results, Figure 3-6 shows the performance evaluation for the out-of-sample projections from the quarterly model assuming perfect foresight of the explanatory variables. The accuracy of the projections is very similar to those of the monthly model, shown in Figure 3-4, despite the simpler specification of the short-run dynamics.

\textsuperscript{10}The results for the quarterly specification using the RBA commodity price index are almost identical.
Table 3.6: Quarterly Dependent Economy Model of Australia.

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<td></td>
<td>(0.528)</td>
<td>(0.528)</td>
<td>(0.388)</td>
<td>(0.388)</td>
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<tr>
<td>$m$</td>
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<td>0.368***</td>
<td>0.573***</td>
<td>0.573***</td>
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<td>(0.141)</td>
<td>(0.134)</td>
<td>(0.134)</td>
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<tr>
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<td>-0.807***</td>
<td>-0.483***</td>
<td>-0.483***</td>
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<tr>
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<td>(0.104)</td>
<td>(0.104)</td>
<td>(0.142)</td>
<td>(0.142)</td>
</tr>
<tr>
<td>$\tilde{p}_M$</td>
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<td>-0.010</td>
<td>(0.186)</td>
<td>(0.186)</td>
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<tr>
<td>$PM$</td>
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<td>-0.628**</td>
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<td>(0.299)</td>
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ERS Test of cointegration

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Short-run dynamics: dependent variable $\Delta s$

<p>| | | | | |</p>
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<td>0.006</td>
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<tr>
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<td>(0.006)</td>
<td>(0.007)</td>
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<tr>
<td>$e_{ct-1}$</td>
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<td>-0.323***</td>
<td>(0.100)</td>
<td>(0.112)</td>
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<td>$e_{ct-2}$</td>
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<td>(0.108)</td>
<td>(0.121)</td>
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<tr>
<td>$\Delta p_x$</td>
<td>-0.383***</td>
<td>-0.397****</td>
<td>(0.135)</td>
<td>(0.147)</td>
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<tr>
<td></td>
<td>-0.328**</td>
<td>-0.374**</td>
<td>(0.143)</td>
<td>(0.154)</td>
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$R^2$ | 0.302 | 0.240 | 0.289 | 0.232 |

***, **** indicate significance at the 1,5,10% level
Figure 3-7: Evaluation of out-of-sample forecasts for the quarterly Australian model.

The expiration dates of commodity contracts limits the construction of the commodity futures index to a 3-month horizon synthetic contract. Commodity prices are assumed to grow at this 3-month rate for each of the quarters in the forecast horizon. This is obviously a gross simplification of market expectations but restrict the analysis to using only data that is observable when the forecast is made. The accuracy of the forecasts generated using the commodity forward premium in the model with the error-correction term lagged two periods is summarised in Figure 3-7.

At horizons of less than 3 quarters the forecasts are no more accurate than the no-change forecast. However, for horizons of 4-6 quarters the forecasts provide a substantial improvement over the no-change forecasts. For longer horizons the improvement diminishes as the assumption of constant commodity price growth becomes increasingly unrealistic. The assessment of the accuracy of the forecasts is virtually identical if the error-correction term is lagged only one period.

Figure 3-8 plots the two-year forecasts made each December from the model that uses constant parameters. The model frequently picks the direction of change in the exchange rate, though typically underestimates the magnitude of change.
Figure 3-8: Forecasts made each December from the constant parameter model using the commodity forward premium.
3.5 Conclusion

The simple dependent economy model tested in this paper provides a good representation of the exchange rates of Australia, Canada and New Zealand. This is especially promising given the poor results of much empirical exchange rate research. The model for New Zealand breaks down at the time of the Asian crisis. This suggests that the model used is an incomplete representation of the exchange rate due to its lack of account for financial factors. It is interesting that given both countries had similar exposure to the Asian crisis that the New Zealand model breaks down at this time, while the Australian model continues to provide a good, though slightly weaker, representation of the Australian dollar. Possibly this is affected by the Reserve Bank of New Zealand's use of a monetary conditions index at this time which induced a simultaneity between monetary policy and the exchange rate, violating an assumption of the estimation strategy. Despite the relatively close correlation of the Canadian commodity price index and export price index, the model of the Canadian dollar is only well specified when the full export price index is used. This makes sense given the much smaller share of commodity exports than for the two smaller countries. While Canada does appear to be a dependent economy, the Canadian dollar's place as a commodity currency is tenuous.

The models for the two larger economies perform well out of sample. For these two countries the model passes the Meese and Rogoff test of outperforming a no-change projection in the out-of-sample period. This is impressive given the out-of-sample period includes the Asian crisis which led to large depreciations in the exchange rates, beyond what many market commentators thought reasonable. The simple model used to construct forecasts of the Australian dollar, which uses only current and past information including the forward premium for commodity futures, shows that there is some predictability of exchange rates, especially at horizons of around one year. Clearly the dependent economy is a special case amongst countries, but it demonstrates there is some promise in developing a more complete understanding of exchange rates.
Bibliography


Appendix – Data

Most data were sourced from the IMF International Financial Statistics (IFS). The country codes for Australia (193), Canada (136) and New Zealand (196), precede the IFS codes given below. Additional data were sourced from the OECD main economic indicators database and Datastream.

Exchange rate – end of period units of national currency per USD, IFS line ..AE.ZF...

Money – Monetary aggregate in national currency divided by real GDP. Real GDP is from IFS line 99BVRZF... and is interpolated to a monthly frequency, money definitions and sources are: Australia – broad money, IFS line 59MC.ZF...; Canada – M2++, sourced from Datastream code CNB1650.; and New Zealand – M3 broad money, IFS line 39MD.ZF...

Unemployment – total non-seasonally-adjusted unemployment rate from the OECD. The New Zealand unemployment rate is only available quarterly and so is interpolated to a monthly frequency.

Commodity price – all commodity price indices are measured in US dollars. The descriptions and sources are: Australia – Reserve Bank of Australia (RBA) Commodity Price Index from RBA Bulletin Table G4; Canada – Bank of Canada Commodity Price Index total and total excluding oil, CANSIM numbers B3300 and B3301; New Zealand – Australia New Zealand Bank commodity price index, source www.anz.com.nz/tools/newslibrary.asp?Commodity_Price_Index

Export price – US dollar export price from IFS line 74..DZF... Note both these series and the Import prices are available for Canada on a monthly basis but for Australia and New Zealand only on a quarterly basis.

Import price – US dollar import price from IFS line 75..DZF...

Proxy import price – US dollar price of exports from industrial countries from IFS line 11074..DZF...

Commodity Futures price index – as described in Chapter 1.