Exchange Rates and Fundamentals: Evidence from Commodity Economies

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Abstract

This paper investigates the empirical disconnect between exchange rates and economic fundamentals that is at the heart of several exchange rate puzzles. In contrast to a literature characterized by notoriously poor in-sample fits and out-of-sample forecast failures, we find that for three major OECD primary commodity producers, nominal exchange rates exhibit a robust response to movements in the world prices of their corresponding commodity exports. Moreover, incorporating commodity export prices into standard exchange rate models can generate a marked improvement in their in-sample performance. In terms of out-of-sample forecast, several commodity-price-augmented specifications offer clear evidence of exchange rate predictability, after accounting for small-sample bias and size distortion in asymptotic tests. However, no single specification is found to provide a consistent forecast improvement over a random walk at *all* horizons and across *all* currency pairs. The empirical approach and evidence shown in this paper can be applied to a much wider set of countries and have particular relevance for monetary policy making and inflation control in small commodity economies.

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1 Introduction

Identifying the relation between economic fundamentals and exchange rates has not been an easy task. Research over the past two decades has repeatedly demonstrated the empirical failures of various structural exchange rate models.[1] When tested against data from major industrialized economies over the floating exchange rate period, canonical exchange rate models produce notoriously poor in-sample estimations, judged by both standard goodness-of-fit criteria and signs of estimated coefficients. In addition, since Meese and Rogoff (1983) first demonstrated that none of the fundamentals-based structural models could reliably outperform a simple random walk in out-of-sample forecasts, numerous research attempts have not been able to convincingly overturn this finding. These empirical challenges led Frankel and Rose (1995) to conclude with doubts about "the value of further time-series modelling of exchange rates at high or medium frequencies using macroeconomic models."

This paper gives standard macro models another chance and investigates whether augmenting them with an additional fundamental can improve their empirical performance.[2] For three OECD commodity-exporting countries (Australia, Canada, and New Zealand), we examine how the world prices of their major commodity exports may influence their nominal currency values. As characterized by MacDonald (1999), research on the various puzzles in the empirical exchange rate literature can be broadly divided into three main areas. The first aims to understand the long-run relationship between nominal exchange rates and various fundamentals. The second area concerns whether fundamentals-based models can produce out-of-sample forecasts that outperform a naive random walk. The third attempts to explain the high volatility and persistence observed in real exchange rate data. Complementing the findings in Chen and Rogoff (2002), this paper focuses on the first two issues concerning nominal exchange rate behavior. We investigate whether world commodity prices can indeed explain nominal exchange rate movements in these economies, and whether augmenting standard models with a commodity price index improves their within-sample fits and out-of-sample forecast performance.

Australia, Canada, and New Zealand are among the many "commodity economies" which have

[1] Frankel and Rose (1995) and Froot and Rogoff (1995) provide comprehensive surveys of the empirical exchange rate literature.

[2] The approach adopted in this paper should be considered complementary to those used in several recent papers that offer promising new directions towards resolving empirical exchange rate puzzles. For example, Kilian and Taylor (2001) develop an exponential smooth threshold autoregressive (ESTAR) model and emphasize the explicit modeling of non-linearity in exchange rate dynamics. Mark and Sul (2001) and Murray and Papell (2002) use new econometric techniques with large panel data setup to improve inference accuracy. Engel (1999), Burstein, Eichenbaum, and Rebello (2002), and Corsetti and Dedola (2002) examine disaggregated data to further understand exchange rate adjustment mechanisms.

a significant portion of their production and exports in primary commodity products.[3] Unlike most other goods where trade may be negotiated bilaterally between countries, commodities are traded in centralized international exchange markets such as the New York Mercantile Exchange (NYMEX), Chicago Mercantile Exchange (CME), and the Chicago Board of Trade (CBOT). This particular market arrangement allows us to observe a clear "world price" for these countries' major exports at any given time. Since most of these countries are not large enough to influence world markets, commodity price fluctuations essentially represent a source of exogenous shocks to their terms of trade.[4] Their heavy reliance on commodity products for export earnings implies that price fluctuations in the world commodity markets would exert pressure on the relative demand for their home currencies, which, in principle, may lead to fluctuations in the values of their floating exchange rates. As most commodity economies either do not have or have not had market-determined floating exchange rates for a sufficiently long period of time, this paper investigates the behavior of nominal exchange rates in Australia, Canada, and New Zealand only.[5] (We note that while we focus on commodity economies where the world prices of their exports are readily observable, in theory any country that exports relatively homogenous products to a highly competitive world market may fit into this analytical framework.[6])

We consider four structural exchange rate models commonly used in the literature – the purchasing power parity (PPP) model, and three variants of the canonical monetary model – and assess their performance in explaining exchange rate movements in the three commodity exporters. We conduct in-sample tests of the models using dynamic OLS, and then augment the models with a country-specific commodity price index.[7] Our results show that for the Australia and New Zealand, commodity prices are a reliable determinant of their exchange rate behavior. An increase in the world price of these countries' commodity exports in general leads to an appreciation of their currencies by roughly the same proportions. In addition, the in-sample fit of several models improves substantially when commodity prices are added into the equations. This suggests that omitted-variable bias may play a role in previous empirical failures. The results for Canada are

[3] In addition to these three countries, Finland, Iceland, Norway, and numerous developing countries, are examples of "commodity economies". In fact, 58 countries are classified as having significant commodity dependency in a recent paper by Cashin, Sespedes, and Sahay (2002).

[4] Chen and Rogoff (2002) provide discussions of and tests for the exogeneity of commodity prices in Australia, Canada, and New Zealand. They also show that world commodity prices better capture the exogenous component of terms of trade shocks than standard measures of terms of trade.

[5] We note that the analysis can easily be extended to more countries, especially with the newly constructed panel dataset of market-based exchange rates in Reinhart and Rogoff (2002).

[6] For example, one could obtain a time series of "world prices" for semiconductor chips and study its effects on the exchange rates of say, Korea and Taiwan. In addition, Sjaastad (1998) applies the commodity-currency approach to the Swiss franc and finds that the short-run movements in the Swiss franc are very well explained by changes in the external prices of Swiss traded goods.

[7] The commodity price indices we use in this paper do not include energy products.

more mixed; in several specifications, a positive commodity price shock leads to a *depreciation* of the Canadian dollar. While the mixed results may be due to Canada's lower commodity dependency compared to those of the Antipodes, we also find evidence that the commodity price-exchange rate connection may not be stable over the thirty year period we examined.

In terms of exchange rate predictability, we use an error-correction framework to investigate both the full-sample predictive content and the out-of-sample point forecast accuracy of the fundamentals-based models. Because these inference procedures are notorious for various small sample biases and size distortions, we employ a bootstrap procedure proposed by Kilian (1999) to examine 1-, 4-, and 8-quarters horizon forecast accuracy. We find that several commodity-price-augmented equations provide strong evidence of exchange rate predictability, showing both statistically and economically significant out-of-sample forecast improvement over a random walk. However, the particular specification that can produce such superior forecast performance appears to depend on the anchor currency, sample period, and the forecast horizon under investigation. This implies that the predictive power may not necessarily be exploitable. While this result may be disappointing, it should not be surprising. Not only is it in accordance with virtually all previous studies on major OECD floating exchange rates over the past two decades, it is also a common conclusion within the broader forecast literature involving other macroeconomic variables.[8]

In addition to offering a new approach by which to analyze long-standing empirical exchange rate puzzles, understanding the impact of commodity price fluctuations on exchange rates may provide important monetary policy lessons for commodity-exporting countries. If commodity price fluctuations indeed induce exchange rate responses, then spot or futures price signals from the world commodity markets may offer relevant information – or even serve as a potential anchor – for the conduct of monetary policy and inflation control.[9] Currently, commodity countries in the world differ widely in their choice of monetary and exchange rate arrangements, which range from inflation targeting under floats to participation in currency unions.[10] While factors such as

[8] For example, Stock and Watson (2001) test the usefulness of 38 potential variables for forecasting US inflation and output. They conclude that "which series predicts what, when, and where is itself difficult to predict: good forecasting performance by an indicator in one period seems to be unrelated to whether it is a useful predictor in a later period."

[9] Incorporating commodity prices into monetary decision making is certainly not a new idea even in the context of non-commodity exporting countries such as the US. Furher (1993) and Woodford (1994), for example, consider using commodity prices as a nonstandard indicator for inflationary pressure in the US. Frankel (2002) further puts forth the proposal that commodity-based economies should peg their currencies to the world price of their commodity exports. While it remains a subject of debate whether commodity prices offer relevant incremental information about future or contemporaneous policy targets in the US, the case may be more clear cut for countries that produce significant primary commodities. A closely related literature asks what the appropriate inflation targets may be in either a closed or open economy setting: headline vs. core inflation, CPI vs. domestic price level inflation. (See, for example, Aoki (2001), Gali and Monacelli (2002), Mankiw and Reis (2002), and Svensson (2000).)

[10] Among the OECD commodity economies, Australia, Canada, and New Zealand have all operated under inflation

industrial structure, capital market conditions, and external debt positions all influence optimal monetary policy strategy, it nevertheless would be useful to understand, *ceteris paribus*, what monetary anchor may be appropriate for economies subject to terms of trade shocks that are as volatile as world commodity prices. Understanding how the appropriate monetary policy may depend on the above-mentioned factors also has important implications for developing commodity economies going through the process of industrialization and liberalization. While a comprehensive analysis of the optimal monetary strategy is beyond the scope of this paper, the empirical evidence presented here clearly invites further structural study of these issues.

The remaining of the paper is organized as the following. Section 2 presents four canonical structural exchange rate models that we will augment with a commodity price index. Section 3 examines within-sample cointegration relation among exchange rates and different fundamentals. Section 4 investigates out-of-sample forecast performance and evaluation. Section 5 concludes.

2 Structural Exchange Rate Models

Three decades of active research since the collapse of the Bretton Woods system has produced numerous structural models of exchange rate determination, linear or non-linear, market microstructure-based or macro fundamentals-based.[11] In this section, we review several classic models of exchange rate determination that have been used prominently in the economic and policy context. The canonical monetary approach to exchange rate modeling typically relies of one or more of the following conditions: long-run purchasing power parity, money market equilibrium, uncovered interest parity, and some treatment of expectations and nominal rigidity. Below we present the PPP model and three variants of the monetary model of exchange rate determination.[12] We then discuss how commodity price shocks influence exchange rate values.

targets for over a decade. Finland, a major exporter of forestry products, participates in the euro area, while Norway, with 70% of its exports in North Sea Oil, operates mostly under a managed exchange rate regime. An even wider range of exchange rate and monetary arrangements can be observed among the developing commodity economies.

[11] See Frankel and Rose (1995) and the forthcoming Journal of International Economics special issue on Empirical Exchange Rate Models for good surveys of the literature.

[12] The monetary model was first introduced by Frenkel (1976) and Mussa (1976). It is often refered to as the workhorse of international finance.

2.1 Canonical Exchange Rate Models:

"Under the skin of any international economist lies a deep-seated belief in some variant of the PPP theory of exchange rates."

- Rudiger Dornbusch and Paul Krugman (1976)

The simplest model of equilibrium exchange rate determination is based on the concepts of purchasing power parity. Whether reflecting the idea that the law of one price holds in traded goods or the Casselian view that the value of a currency should reflect its relative purchasing power, the concept of PPP is central to international economics.[13] The first model we consider thus tests how exchange rates in commodity economies relates to their relative CPI price level. Recognizing the existence of transport costs, risk premia, and other factors, we allow for deviations from "absolute PPP" and consider the following specification:

$$s_t = \alpha + p_t^* - p_t + \varepsilon_t$$

Here all variables are expressed in logarithmic terms. s_t is the foreign currency price of a unit of home (commodity) currency so a larger number means an appreciation of the home currency. p and $p^*$are home and foreign CPI-price levels respectively, and ε_t represents a stationary disturbance here as well as in all equations below.

We next consider two variants of the flexible price monetary model. The monetary models build upon PPP by imposing additional structural restrictions, treating non-money assets as perfect substitutes. These additional restrictions, if true, should lead to better asymptotic performance.[14] One such additional assumption is money market equilibrium, which states that the log of real money demand depends linearly on the log of real income and the nominal interest rate:

$$m_t - p_t = \beta_y y_t - \beta_i i_t + \varepsilon_t$$

Assuming an analogous equation for the foreign country and equalized income elasticities and interest semi-elasticities of money demand across countries, the price levels in the relative PPP equation above can be substituted out. The exchange rate then becomes a positive function of the relative money stocks and the nominal interest differentials between the two countries, and a negative function of their relative real income. This form of the monetary equations is considered

[13] In this paper we focus on the Casellian PPP and consider CPI-based PPP. PPP based on Samuelson view of no goods market arbitrade would favor a PPI-based PPP.

[14] Indeed, Mark and Sul (2001) provide some evidence that the linkage between exchange rates and monetary fundamentals is tighter than that between exchange rates and purchasing power parities.

in Bilson (1978), Frenkel (1976), and more recently in MacDonald and Taylor (1994) and Flood and Rose (1995). Nominal interest rate differentials here reflect inflation premia. Facing an increase in the expected inflation at home, agents would switch from domestic currency into bonds, inducing a depreciation of the home currency.[15]

$$s_t = \alpha + (m_t^* - m_t) - \beta_y(y_t^* - y_t) + \beta_i(i_t^* - i_t) + \varepsilon_t$$

We next consider the asset approach which further incorporates expectational effects and asset market linkages in exchange rate modeling. Assuming that domestic and foreign assets are perfect substitutes (it would be straight forward to incorporate exogenous risk premia), international capital market equilibrium is then given by the uncovered interest parity (UIP) condition:

$$i_t^* - i_t = E_t(s_{t+1} - s_t)$$

Incorporating the UIP into the flexible price monetary model above gives us a first-order stochastic difference equation for the nominal exchange rate, whose current value embodies its expected future values (and thus future development of fundamentals). Solving for the rational expectation no-bubbles solution, we can express exchange rate as the expected present-value of relative money stock and relative real income, a model emphasized in several important recent papers such as Mark (1995) and Kilian (1999), among others. It is then common to assume the two sets of fundamentals to follow a driftless random walk, leading to the following reduced-form equation:

$$s_t = \alpha + (m_t^* - m_t) - \beta_y(y_t^* - y_t) + \varepsilon_t$$

Finally, we consider a monetary model that incorporates short-term price rigidities, following the work of Dornbusch (1976) and Frankel (1979). In the presence of short-run price stickiness, purchasing power parity condition would be violated temporarily, and the relation between interest rates and the exchange rate needs to capture the short-term liquidity effects of monetary policy. Frankel (1979) proposes using a real interest rate differential term which consists of a nominal interest rate with short-term maturity, capturing the liquidity effects, and an inflation expectation component that operates through Cagan money demand function. The reduced form equation can be represented as follows.[16] (We note that in this model, interest differentials enter the exchange equation with the opposite sign as in the flexible price model presented above.)

$$s_t = a + (m_t^* - m_t) - \beta_y(y_t^* - y_t) - \beta_i(i_t^* - i_t) + \beta_\pi(\pi_t^* - \pi_t) + \varepsilon_t$$

[15] Long-term interest rates may be more appropriate here to capture inflation premia, but in the empirical analysis, 3 month T-bill rates are used for all interest rates.

[16] Cheung, Chinn, and Pascual (2002) considered this equation as the representative of the 1970's vintage models.

The above models along with other variants have all been tested extensively without much empirical success.[17] In this paper, we focus on the following four specifications that relate exchange rates s_t linearly to a set of fundamentals f_t:[18]

$$s_t = \alpha + f_t + \varepsilon_t \tag{1}$$

where f_t is model-dependent and are described by the equations below:

Relative PPP Model:

$$f_t = \beta_p(p_t^* - p_t) \tag{2}$$

Asset Approach Flexible Price Monetary Model:

$$f_t = \beta_m(m_t^* - m_t) - \beta_y(y_t^* - y_t) \tag{3}$$

Flexible Price Monetary Model:

$$f_t = \beta_m(m_t^* - m_t) - \beta_y(y_t^* - y_t) + \beta_i(i_t^* - i_t) \tag{4}$$

Sticky Price Monetary Model:

$$f_t = \beta_m(m_t^* - m_t) - \beta_y(y_t^* - y_t) - \beta_i(i_t^* - i_t) + \beta_\pi(\pi_t^* - \pi_t) \tag{5}$$

The regression coefficients have the following interpretations and theoretical values: β_p, the coefficient on the relative CPIs, and β_m, the elasticity with respect to money stock, should be unity. β_y represents the income elasticity of money demand, and β_i and $\beta_\pi$the interest and expected inflation semi-elasticity.

2.2 Commodity Price Shocks and Exchange Rates

Commodity price shocks, or more generally terms of trade fluctuations, may affect the value of the exporting country's currency through multiple channels. In this section, we present a brief discussion on this topic in order to justify why commodity prices may be an important additional control in empirical exchange rate equations. We note that our goal is not to advocate or test a particular channel through which they may affect exchange rates.

First, consider the small open economy model with non-traded goods presented in Chen and Rogoff (2002). It shows that an increase in the world price of a country's commodity exports will exert upward pressure on its real exchange rate, through its effect on wages and the demand for

[17] See Frankel and Rose (1995) for examples of other augmentations that have been proposed and tested, such as models which incorporate current account information, productivity differences and so on.

[18] Following Cheung et all (2002), we also tested a Balassa-Samuelson productivity-augmented model specification. The model does not work well (results available upon request).

non-traded goods (a channel similar to the standard Balassa-Samuelson effect). In the presence of nominal price rigidities, however, the exchange rate, instead of prices, will have to do the adjustment to preserve efficient resource allocation. For example, when the price of non-tradable goods are sticky and unable to respond to the upward pressure induced by a positive terms of trade shock, exchange rate would need to appreciate to restore the efficient relative price between traded and non-traded goods.

The portfolio-balance model provides another justification for the relevance of commodity prices in exchange rate determination. This class of models treats domestic and foreign assets as imperfect substitutes, thus exchange rates are dependent on the supply and demand for all foreign and domestic assets, not just money. For an economy that relies heavily on commodities for export earnings, a boom in the world commodity market would typically lead to a balance-of-payments surplus and an accumulation of foreign reserves, exerting pressure on the relative demand of their home currencies. This would then lead to an appreciation of the domestic currency.[19]

In reality, these channels may be at work simultaneously, justifying including a good measure of terms of trade into exchange rate analyses. As discussed and investigated more fully in Chen and Rogoff (2002), identifications off of traditional terms of trade measures – the relative price index of exports to imports – are mostly complicated by price stickiness and potential mechanical correlations, endogenous pricing behavior, and other factors. Commodity economies, with its particular global trading arrangement, represent a good experiment to test out the terms-of-trade hypothesis.

We consider four commodity price-augmented equations listed below:

Augmented Relative PPP Model:

$$s_t = \alpha + \beta_{cp} p_t^{com} + \beta_p (p_t^* - p_t) + \varepsilon_t \tag{6}$$

Augmented Asset Approach Flexible Price Monetary Model:

$$s_t = \alpha + \beta_{cp} p_t^{com} + \beta_m (m_t^* - m_t) - \beta_y (y_t^* - y_t) + \varepsilon_t \tag{7}$$

Augmented Flexible Price Monetary Model:

$$s_t = \alpha + \beta_{cp} p_t^{com} + \beta_m (m_t^* - m_t) - \beta_y (y_t^* - y_t) + \beta_i (i_t^* - i_t) + \varepsilon_t \tag{8}$$

Augmented Sticky Price Monetary Model:

$$s_t = \alpha + \beta_{cp} p_t^{com} + \beta_m (m_t^* - m_t) - \beta_y (y_t^* - y_t) - \beta_i (i_t^* - i_t) + \beta_\pi (\pi_t^* - \pi_t) + \varepsilon_t \tag{9}$$

where p^{com} represents the world price in US dollar of a country's major commodity exports.[20]

[19] Studies on many developing countries (Indonesia, Kenya, Columbia..etc. Edwards (1986).

[20] We consider non-energy commodity prices only. None of these countries is an obvious large net exporter of

β_{cp} should enter with a positive sign according to the channels discussed above. The remainder of the paper will be empirical in nature, aiming to examine the following questions: 1) Does the world price of these countries' commodity exports explain movements in their exchange rates? 2) Controlling for any such comovement, do standard models fit the data better? and 3) Do the commodity price augmented models offer more predictive content than standard models and/or the random walk? We thus remain agnostic about the implications of any particular theory discussed above.

3 Empirical Exchange Rate Determination

3.1 Background and Data

Empirical exchange rate puzzles concern mainly the behavior of floating exchange rates between countries with open trade and liberalized capital markets, where the currency values are most likely to reflect various macroeconomic market forces. For this reason, we focus on OECD commodity economies with sufficiently long floating exchange rate history to allow meaningful time series analysis. Australia, Canada, and New Zealand are good examples of such open economies with little market frictions and government interventions. In terms of commodity dependency, Australia and New Zealand both have over half of their exports in various primary commodity products, and in Canada, the share in recent years is roughly a quarter. (We refer interested readers to Chen-Rogoff (2002) and references there within for more detailed descriptions of these economies and their commodity exports.)

We look at the values of the Australian, Canadian, and New Zealand dollars relative to three anchor currencies: the US dollar, the British pound, and the Japanese Yen. Figures 1 through 3 show the three countries' end-of-period quarterly nominal exchange rates, their Consumer Price Index relative to that of the anchor country, and also the nominal US dollar price index of their respective primary commodity exports. We observe that relative consumer prices capture little of the short- to medium-term movements in the exchange rates, though appear to share the same longer run trend with exchange rates. Commodity prices on the other hand appear to capture more of the higher frequency swings in the exchange rates. In the following subsections, we test for and establish the existence of cointegration relations amongst the series, and then using the dynamic OLS (DOLS) -3procedure of Stock and Watson (1993) to estimate the cointegrating vectors.

In terms of data, we follow the literature and use M1 money as the measure for money stock, except in the case of the UK, where M0 is used. Money, CPIs, real output, and the rate of returns

energy commodities as they are with non-energy products.

on three-month Treasury bills are all taken from the International Financial Statistics of the IMF (see Data Appendix for more details).

3.2 Integration and Cointegration

The non-stationary behavior of exchange rates and prices raises concern about spurious regressions. However, theories such as PPP and the monetary models described above impose theoretical restrictions on the relative movements of these variables, such that deviation of exchange rates from the various fundamentals should be stationary. In this section, we test for the degree of integration of the series and provide evidence that the above models can be viewed as cointegrating relationships among non-stationary variables. That is, the series are cointegrated in the Engle and Granger (1987) sense, and the cointegrating equation may be interpreted as a long-run equilibrium relationship among the variables.[21]

Representative unit root test results or nominal exchange rates, consumer prices, commodity prices, money stocks, and real output for the relevant countries are reported in Appendix Table A. Unit root tests are conducted using procedures proposed by Elliot, Rothenberg, and Stock (1996) and by Kwiatkowski, Phillips, Schmidt, and Shin (1992). The results confirm the order of integration in these series to be one. We then considered two tests for cointegration relations. The first is the two-stage Engle-Granger Augmented Dickey-Fuller (ADF) procedure. The ADF tests for the stationarity of the OLS residuals from the exchange rates-fundamentals regressions. We report its results in Tables 1 to 3 along with the DOLS estimation results. The second approach, reported in the appendix, is the multivariate cointegration test procedure proposed by Johansen (1988). It detects if any linear combination of the variables in our models is stationary. Although in general such cointegration tests suffer from the problem of low power, we do observe some supporting evidence of cointegration among most variables.[22] As the Johansen procedure is well-known to produce widely dispersed estimates, we proceed with dynamic OLS in the next section to estimate the cointegrating vectors.

3.3 Dynamic OLS Estimations

For I(1) series with one cointegration relation, the dynamic OLS estimation procedure proposed by Stock and Watson (1993) produces efficient estimates of the cointegrating vector. This procedure involves augmenting OLS regressions with lead and lag values of the first differences of each

[21] Cointegration refers to a linear combination of nonstationary variables. Of course, it is possible that a nonlinear long-run relationship exists among the set of integrated variables.

[22] We note that a couple newer tests for cointegration may be appropriate here, such as the procedure proposed by Horvath and Watson (1996), and the point optimal tests discussed in Jansson (2002).

regressors. That is, we estimate the cointegrating vector β between the exchange rate s_t and a set of fundamentals f_t using the following DOLS specification:

$$s_t = a + \beta f_t + \sum_{j=-p}^{p} \delta_j \Delta f_{t-j} + \varepsilon_t \tag{10}$$

where fundamentals f_t, as before, will be model dependent and are described in equations $(2)-(5)$ and $(7)-(10)$. Table 1 below shows results pre- and post- inclusion of commodity prices, using US dollar as the anchor currency. Estimation results for the three commodity currencies relative to the British pound and the Japanese Yen are included in Tables 2 - 3.

From the coefficient estimates in Table 1, we first note that for Australia and New Zealand, the US dollar price of their respective commodity exports appear to be a consistent factor in explaining their exchange rate movements. For the Australian-US exchange rate, a one percent increase in the price of Australian commodity exports appreciates its value by about 0.8 percent on average. For New Zealand, the exchange rate responses are larger, at around 1.4 percent. These estimates appear quite robust to using other base currencies, as shown in Table 2 and 3. The exchange rate-commodity price connection for Canada, however, appears weaker and less consistent, even regarding the direction of the response. In the majority of the regressions, a one percent increase in the world price of Canada's commodity exports actually leads to a with an elasticity coefficient of around 0.3 percent *depreciation* of the Canadian dollar. Judging from additional regression results (not reported here), the Canadian coefficient estimates appear to be sensitive to the sample period considered. Using data starting in 1980, for example, we see commodity price elasticity estimates comparable to those for Australia or New Zealand. Further analysis of parameter stability and possible regime switches over the 30 year period is in order to understand better the Canadian dollar-commodity price connection.

Turning to the standard models, we see that without the commodity price term, the models indeed perform extremely poorly, both in their overall fit and in producing coefficient estimates with sensible signs. This is especially apparent when the anchor currency is either the US dollar or the British pound. (The models in general seem to perform much better in explaining the Yen cross rates.) Note that the inclusion of commodity prices not only substantially improves the fit of the models in most cases, it also leads to estimates with more sensible signs (though they are not always significant) in some specifications.[23] The PPP equation for the Australian-US exchange rate stands out as an good example of how an otherwise extremely poor-fitting estimation becomes one that fits the theoretical prediction well after the introduction of the commodity price term. From

[23] We note that relative short-term interest rate differentials rarely follow the signs predicted by the structural models, with or without the inclusion of commodity prices.

the two-stage Engle and Granger tests, we also see that the inclusion of commodity prices can turn otherwise nonstationary residuals from OLS regressions to one supporting cointegration. These results provide support for the view that at least for commodity economies, terms of trade shocks, as captured by commodity price movements, are important for explaining exchange rate behavior, and that their omission may explain some of the earlier failures of these empirical exchange rate equations.

4 Exchange Rate Predictability

The most robust conclusion from two decades of post-Meese-Rogoff (1983) research is that its original finding is remarkably resilient. While several subsequent research claim to have found models with forecast performance superior to that of a random walk (especially at long-horizons), the robustness of their results and the inference procedures have been called into questions. (See Berben and van Dijk (1998), Berkowitz and Giorgianni (2001), Groen (1999), Kilian (1999), and Rossi (2002), among others, for a good coverage of the debate.) The relatively short periods of floating exchange rate regimes and the non-stationarity in the data present thorny inference difficulties for establishing reliable results. A few recent papers have shown some success in outperforming a random walk forecast, but their approaches mostly depart from the traditional linear time series analysis.[24] Even under these alternative setups, they show that while possible, it is indeed difficult to beat a random walk forecast using fundamentals-based specifications. In this section, we return to the conventional linear time series setup for analyzing exchange rate predictability.[25] By using commodity prices as an additional fundamental, we evaluate the performance of standard models in providing forecast information and ask whether the commodity-price-augmented models can outperform a random walk in short- to medium-run horizon forecasts.

A principal feature of cointegrated variables is that their time paths are influenced by deviations from the long-run equilibrium. As such they may be represented in a dynamic error correction framework. In this section, we examine the usefulness of macroeconomic fundamentals in providing exchange rate forecasts, adopting an error-correction specification. We first conduct in-sample

[24] For example, Kilian and Taylor (2001) developed a non-linear exponential smooth threshold autoregressive model that outperforms the random walk model, using relative prices as fundamentals. Mark and Sul (2001) use a panel cointegration framework to look at exchange rate predictability in 19 OECD countries and find some predictability in the short-run. Faust, Rogers, and Wright (2001) focus on using real time data, and show that using revised fundamentals in conventional empirical exchange rate equations might explain their poor performance.

[25] Pooling countries into a panel setup can indeed increase the statistical power of tests and overcome part of the inference difficulties. However, it is unclear that a panel approach would necessarily draw the relevant variation, as currency values in different countries may be driven by very different forces (such as monetary policy and exchange rate regimes) and have different underlying data generating processes. Rapach and Wohar (2001) and Neely and Sarno (2002) discuss this issue and cast doubt on the validity of pooling data across countries.

tests on whether deviations of the current exchange rate from its "fundamental value" suggested by the structural models can help predict future exchange rates at the 1-, 4-, and 8-quarter horizons.[26] We next look at simulated out-of-sample point forecasts based on these models, and compare them to the results from a random walk. All estimations and forecasts are conducted with and without the inclusion of commodity prices, in order to assess whether this additional "fundamental" affects the performance of standard models. The section below addresses the multitudes of inference complications involved in these estimations. The next two sections present some illustrative results, without rigorously addressing the numerous the potential biases and inference complications involved in these analyses. This is in part due to the fact that the correct inference procedure is often model- and/or data-dependent (or perhaps unknown to us.)[27] In the final subsection, we consider the bootstrap procedure proposed first by Mark (1995) and modified by Kilian (1999), and evaluate the statistical significance of the performance of a few models, using tabulated small-sample critical values. We demonstrate that the illustrative results hold and that some commodity price augmented models do lead to both statistically and economically significant forecast improvement over a random walk model.

4.1 Error Correction Framework and Complications

If exchange rates are cointegrated in the long-run with the fundamentals as suggested by the structural models, short-term deviations of the current exchange rate from its "fundamental value" should help predict future exchange rate movements.[28] In the error correction specification below, exchange rate predictability would be indicated by a positive and significant full-sample slope coefficient estimate λ_k.

$$\triangle s_{t+k} = s_{t+k} - s_t = \alpha_k + \lambda_k[\beta f_t - s_t] + \upsilon_t \tag{11}$$

where f_t is a vector of fundamentals as described by the models described in section 2 (see (2) $-$ (5) and (6) $-$ (9)). β is the cointegrating vector between the set of fundamentals and the exchange rates; k is the forecast horizon, and υ_t a stationary disturbance. The equation says that when the current exchange rate falls below its long-run value implied by the fundamentals, an appreciation should occur in the future. A one-sided test of the hull hypothesis of $H_0 : \lambda_k = 0$

[26] These in-sample predictive regressions test whether the slope coefficient on the error correction term is zero.

[27] Evaluating inflation and output forecast performance of different variables, Stock and Watson (2001) state "due to the current lack of knowledge of the finite sample performance of these procedures, we do not report standard errors of forecast equality." The subsections below provide further discussion.

[28] This is assuming that exchange rates, rather than the fundamentals, do the "correction." This is a reasonable assumption given the other fundamentals we consider.

against the alternative hypothesis of $H_1 : \lambda_k > 0$ is thus a test of the predictability of exchange rate at k horizons in the future. While the idea is straight forward, several thorny inference difficulties are involved in these estimations. Below is a brief discussion of a couple of the complications commonly discussed in the literature.

First, the error-correction framework assumes the existence of a known cointegration vector β between the exchange rate and the fundamentals that would ensure the stationarity of the error correction term, $(\beta f_t - s_t)$. In practice, however, it is difficult to know what β should be, especially given the short sample sizes where cointegration tests and estimations have low power. The problem is further complicated by the fact that the asymptotic null distribution of test statistics for λ_k depends on whether the error-correction term is stationary or not, as discussed in Cavanagh, Elliott, and Stock (1996). Even if the error-correction term is indeed stationary, conventional inference procedure on λ_k again breaks down in finite samples if the error correction term is highly persistent (see Campbell and Yogo (2002) in the context of stock return predictability).[29]

In this paper, we consider two alternative approaches to the problem of unknown β, recognizing that neither is ideal. We first adopt a common approach in the literature which pre-imposes the theoretical values of β implied by the structural models before running the error-correction equation (see Chinn and Meese (1995), Mark (1995), Kilian (1999), among others). While this approach is particularly justifiable for testing theory/hypothesis-testing, it is less helpful in analyzing the predictive content of exchange rates, given how poorly theoretical parameter values are supported by actual data. And as discussed more fully in Berkowitz and Giorgianni (1999), imposing them may result in non-stationary regressors and lead to bias in favor of finding predictability. For commodity economies, we hope that these issues are mitigated, given the results we see in Section 3 that the evidence for cointegration appears stronger with the inclusion of commodity prices into the standard models. An alternative approach, adopted by MacDonald and Taylor (1994) among others, is to pre-impose the full-sample estimated cointegration vectors in the error correction regressions. We use the coefficients obtained from the DOLS regressions in the previous section for β. While this approach introduces parameter estimation uncertainty, it may better capture the empirical underlying cointegration relations between the fundamentals and the exchange rates.

Another inference difficulty is that when the forecast horizon is large than one, the predictive equation (11) above involves overlapping observations. In small samples where the overlap periods is large relative to the sample size, test statistics such as the in-sample R^2 and the t-statistics are non-standard, and the coefficient estimate $\widehat{\lambda}_k$ is biased away from zero due to size distortions.[30] To

[29]The two-step efficient testing procedure proposed in Campbell and Yogo (2002) may be an alternative and complementary approach here.

[30]This bias in favor of finding predictability as forecast horizon increases is well documented in studies of stock

address this problem, we adopt the bootstrap procedure proposed first by Mark (1995) and modified by Kilian (1999), and evaluate the statistical significance of the model's forecast performance based on the tabulated small-sample critical values.

4.2 In-Sample Predictive Regressions

Equation (11) provides the basis for full-sample test of exchange rate predictability:

$$\triangle s_{t+k} = \alpha_k + \lambda_k[\beta f_t - s_t] + \upsilon_t \tag{11}$$

We conduct a one-sided test of the hull hypothesis of $H_0 : \lambda_k = 0$ against the alternative hypothesis of $H_1 : \lambda_k > 0$. As discussed above, we consider two alternative sets of values for the cointegrating vector β, trading off parameter uncertainty and potential mis-specification. For the theoretical values of β, we follow McCracken and Sapp (2002), among others, and use: $\beta_p = 1; \beta_m = 1; \beta_y = -1; \beta_\pi = 1$;and as discussed in section 2, $\beta_i = 1$ in the Flexible Price Monetary Model and $\beta_i = -1$ in the Sticky Price Monetary Model. For commodity price elasticities, we take the empirical values from the full-sample estimations presented above: $\beta_{cp}^{AUS} = 0.8; \beta_{cp}^{CAN} = -0.3; \beta_{cp}^{NZL} = 1.4$. In addition, we conducted robustness checks on these parameter assumptions, using values proposed in Meese and Chinn (1995).[31]

Table 4 presents the estimated values of λ and their significance level for the four "theoretically-constrained" predictive equations based on the structural models discussed earlier. Predictive regressions are run at both the 1- and 4-period horizons for Australia and New Zealand, and for Canada, 1-, 4-, and 8-quarter forecasts are conducted. While these results are illustrative, we do observe that in some specifications (such as the PPP model in Australia and New Zealand), the inclusion of commodity prices turns slope coefficients that are otherwise insignificantly different from zero into larger, and significantly positive ones. We also note that despite using a naive correction for serial correlation – the Newey-West standard errors – we still observe more evidence of predictability as forecast horizons lengthen. This likely reflects the bias due to overlapping observations we discussed earlier.[32] For 1-quarter ahead forecast using Australian-US exchange rate, the inclusion of commodity prices in general improves the forecast outcome; however, except under the PPP specification, the improvements do not appear economically significant. For Canada, these equations perform dismally, even in 8 quarters ahead (where there is more bias towards finding

returns such as Fama and French (1988), Campbell and Shiller (1988), and in the exchange rate literature by Mark and Sul (2001), Kilian (1999), Berkowitz and Giorgianni (1999), among others.

[31] We reach the same qualitative conclusions using alternative parameter values. Results are available upon request.

[32] As we know that the Newey-West method, a non-parametric approach, does not work well in small samples.

predictability). The inclusion of commodity prices does not seem to make a big difference for the New Zealand dollar (again, except in the PPP regressions), and we observe evidence of exchange rate predictability from all three versions of the monetary models at both 1 quarter and 4-quarters ahead.

Next we turn to examine an alternative set of β values. We use the estimated cointegration coefficients obtained from DOLS in Table 1 through 3. Imposing the in-sample estimated β values in general suggests a stronger role for commodity prices and offers more support for exchange rate forecastability in Australia and New Zealand. The slope coefficients are larger, and the associated t-statistics are higher as well. For Canada, the results again do not provide strong support for the view of exchange rate predictability, with or without the help of commodity prices.

We reconsider a few of these specifications in Section 4.4 and evaluate their statistical significance using the Kilian (1999) bootstrap procedure.

4.3 Simulated Out-of-Sample Forecasts

In this section, we follow the tradition of the Meese and Rogoff (1983) literature to consider whether fundamentals-based models can outperform a random in out-of-sample forecast accuracy. We again base our analysis on error correction dynamics and consider forecast equation of the following form:

$$\triangle s_{t+k} = \alpha_k + \lambda_k[\beta_t f_t - s_t] + \upsilon_t \tag{12}$$

where f_t is a vector of fundamentals as before, and β_t is the cointegrating vector between the fundamentals and the exchange rates. (We note that unlike equation (11) above, here β_t is presented with a time t subscript, as will be explained later.) We use a rolling regression procedure to generate out-of-sample exchange rate point forecasts from the PPP and the monetary models, and compare the results to the forecast outcome from a random walk (without drift).

Again, we consider two alternative sets of values for β_t. The first, as in previous section, is to constrain β to be the values implied by the relevant theories. That is, in the PPP model, exchange rates and relative prices move one-for-one. In the monetary models, the coefficient on relative money stock and the income elasticity of money demand are set to 1 and -1 respectively. Interest semi-elasticity is set to 1 in the Flexible Price Monetary Model and -1 in the Sticky Price Monetary Model. Inflation semi-elasticity is taken as 1. For commodity price elasticities, we take the empirical values from the full-sample estimations presented above: $\beta_{cp}^{AUS} = 0.8; \beta_{cp}^{CAN} = -0.3; \beta_{cp}^{NZL} = 1.4$.[33]

[33] Robustness tests are conducted with alternative coefficient values, such as an interest differentials semi-elasticity

The second forecast approach does not impose any β, but rather allows the coefficients to be determined by the data simultaneously as we go through the recursive procedure described in the next section (hence the subscript t in β_t). This approach is in spirit closer to the original Meese-Rogoff rolling regression approach, and rather than hypothesis or theory testing, its focus is more on the predictive content or forecastability of exchange rates using macro fundamentals. The additional uncertainty and potential errors associated with parameter estimations may deteriorate forecast accuracy, but to the extent that the true cointegrating vector may differ from its theoretical predictions, the data may pick up this difference and produce more accurate forecast.

We use equation (12), the appropriate $\beta's$, and the relevant fundamentals associated with each of the four structural models to generate short horizon forecasts. Again, we evaluate each model with and without including the commodity prices.

4.3.1 Recursive Out-of-Sample Forecast Procedure

For each of the 8 specifications described above (4 models with 2 sets of $\beta's$), out-of-sample forecasts are generated for three different forecast horizons: $k = 1$, 4, and 8 quarters. We also look at forecast performance for three different evaluation periods: the last $20, 30$, and 40 quarters.

Following Mark (1995), Faust, Rogers, and Wright (2001), among others, we conduct forecast exercise using the following procedure: In-sample OLS estimates are obtained for equation (12) using observations available through $1991Q3$, $1994Q1$, and $1996Q3$, the three "end dates". (These three dates correspond to using the last 40, 30, and 20 quarters of sample data for forecast evaluation.) The $k = 1$ regression result is then used to forecast exchange rate return for $1991Q4$, $1994Q2$, and $1996Q4$, and similarly for $k = 4$ and 8. To generate the next forecast, the estimation sample is updated by one period, adding the observation for $1991Q4/1994Q2/1996Q4$ into the OLS estimations. The procedure is repeated until we reach the end of the sample at $2001Q2$.

Forecast outcome is produced using the PPP model and the three monetary models, with and without coefficient constraints as explained above, and with and without the commodity price term in the estimation.[34] We focus on the root mean squared errors (RMSEs) generated by the model over the simulated out-of-sample sub-period, relative to the RMSEs of the benchmark random walk without drift computed over the same period.[35]

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of 0.5, following Faust, Rogers, and Wright (2001).

[34] For each country pairs, there are 2×72 forecast results: each of the 4 models is estimated *twice* with and without commodity prices for 3 forecast horizons and for 3 forecast evaluation periods. Each specification is done *twice* with and without coefficient constraints, which are reported separately.

[35] The consensus in the literature is that a random walk without drift outperforms a random walk with drift in forecasting, and hence one without drift is a harder model to beat. Kilian (1999), on the other hand, argues that random walk with drift may be the more appropriate benchmark.

As discussed earlier, testing for relative forecast accuracy in small samples involves numerous complications, and the proper statistics and inference procedures for comparing forecast accuracy are very much sensitive to the models under comparison, the forecast specifications, the sample sizes, and the underlying time series properties of the variables under consideration.[36] For example, a commonly adopted test statistics, the Diebold-Mariano (1995) statistics, are useful for drawing inference about predictions for non-nested models when the prediction process does not involve regression estimates. When the models are non-nested, heteroskedasticity and autocorrelation consistent (HAC) standard errors for this relative RMSEs can be computed following West (1996). Clark and McCracken (2001) consider asymptotic distribution for nested models. However, given our small sample size, it is unclear that these procedures apply, as stated in Stock and Watson (2001). As such, the results in figure 4 and 5 are for illustrative purposes only.

Charts in Figure 4 show representative results for specifications estimated with coefficients constrained to their theoretical values. We note that the inclusion of commodity prices overall do not appear to provide much additional forecast improvements to the standard models, except in a few instances such as the PPP model for the New Zealand dollar-Yen rate. Turning to Figure 5 where the cointegrating vector is estimated within-sample, we see very different results. In general, incorporating commodity prices *does* improve forecast performance, and the improvement becomes more pronounced as forecast horizons increase. In addition, at least for results using US dollar as an anchor, the pattern seems to confirm the findings in Mark and Sul (2001): that the linkage between the exchange rate and monetary fundamentals is tighter than that between the exchange rate and PPP. Comparing the results in figures 4 and 5 suggest that while the structural models may not perform well, the fundamentals themselves nevertheless may contain useful information for forecasting future exchange rate movements.

4.4 The Bootstrap Approach

Long-horizon regression tests, as discussed above, can bias in favor of finding predictability due to severe size distortions, spurious regression fits, and small-sample bias in the estimates of regression coefficients and asymptotic standard errors. To mitigate the bias in asymptotic critical values of the test statistics, Kilian (1999) builds upon the bootstrap approach proposed by Mark (1995) and calculates critical values based on bootstrap approximations of the finite sample distribution of these test statistics. As explained in Kilian (1999), these critical values, based on the percentiles

[36] See Clark and McCracken (2001) for tests of equal forecast accuracy for nested linear models. Corradi, Swanson, and Olivetti (2001) outline conditions under which the Diebold-Mariano forecast comparison test can be applied to models of cointegrated systems. Clements and Hendry (2001) highlight the fact that deterministic shifts often lead to systematic forecast failure.

of the bootstrap distributions, automatically adjust for the increase in the dispersion of the finite-sample distribution of the test statistics that occurs in near-spurious regressions as the sample size grows. So the bootstrap inference is immune to the near-spurious regression problems in Berkowitz and Giorgianni (1999).

We adopt this approach to examine a few representative models where the cointegrating vector is imposed according to theoretical predictions. We look at three separate statistics to evaluate the statistical significance of the model's forecast performance: the slope coefficient λ_k, the RMSE ratios, and the Diebold-Mariano (1995) statistics. Table 5 presents the test statistics with their associated p-values. We see that these specifications indeed provide strong support for exchange rate predictability and offer significant outperformance over a random walk. However, we also note that the particular specification that can produce such superior forecast performance appears to depend on the anchor currency, sample period, and the forecast horizon under investigation. This implies that the predictive power may not necessarily be exploitable. While this result may be disappointing, it should not be surprising. Not only is it in accordance with virtually all previous studies on major OECD floating exchange rates over the past two decades, it is also a common conclusion within the broader forecast literature involving other macroeconomic variables, as summarized in Stock and Watson (2001): "which series predicts what, when, and where is itself difficult to predict: good forecasting performance by an indicator in one period seems to be unrelated to whether it is a useful predictor in a later period."

5 Conclusion

Standard macro fundamentals-based models of exchange rate determinations are well known to offer little empirical value in explaining exchange rate behavior in OECD countries. This paper presents evidence that at least for some of these countries where an exogenous terms-of-trade variable can properly be identified, a good portion of their exchange rate movements can be explained. For Australia and New Zealand, the world price of their commodity exports appears to have a stable and consistent impact on the value of their currencies over their floating rate periods. A rise in the commodity price index is associated with an appreciation of the currency by roughly the same proportion (around 0.8 for Australia and 1.4 for New Zealand). The evidence for Canada is more mixed, and the correlation between commodity price and nominal exchange rate tends to show up as negative, with an elasticity coefficient of around -0.3. The mixed results for the Canadian dollar is likely due to its much lower commodity export dependency (at around a quarter of its total exports), compared to the cases of the Antipodes, where various commodity products represent around half of their exports.

Using commodity prices as an additional fundamental, we re-examine the performance of four standard macroeconomic models in explaining both in- and out-of-sample nominal exchange rate behavior. We find that the inclusion of commodity prices improves the in-sample fit of several models, and in general offers more support for long-run cointegration relations between exchange rates and fundamentals. These findings suggest that properly accounting for terms-of-trade fluctuations may be an important piece of the puzzle for explaining previous empirical failures. In terms of the predictive content of these models in short- to medium- horizon forecasts (1 quarter to 2 years), this paper shows that the inclusion of commodity prices as an additional fundamental can improve the predictive accuracy of some, though not all, of the standard models. Using a bootstrap procedure proposed by Kilian (1999) to account for various biases and size distortions involved in these estimations, we find that several commodity-price-augmented equations not only provide strong evidence of exchange rate predictability, they also outperform a random walk in forecast accuracy by a statistically and economically significant amount. However, there does not appear to be one single model that can provide such superior predictive performance in *all* forecast horizons and across *all* country pairs.

Although these positive empirical findings may not be immediately applicable for explaining-ing the behavior of other major OECD currencies, they nevertheless offer important lessons for developing commodity economies as they go through the process of industrialization and liberalization. To the extent that world commodity prices induce exchange rate fluctuations, spot or futures price signals in the world commodity markets may offer additional information for monetary policy making and inflation control in these economies.

A Appendix: Unit Root and Cointegration Tests

To ascertain the degree of integration in the individual series, two unit root tests are performed on both the levels and the first differences of all relevant variables. First, the general least square-based augmented Dicky Fuller test (ADF-GLS) developed by Elliott, Rothenberg, and Stock (1996) is performed to test the null hypothesis of an autoregressive unit root. It is well known that standard unit root tests possess low power against the alternative of a stationary but highly persistent series. While there is no strictly uniformly most powerful invariant test for the unit root hypothesis, this modified Dickey Fuller test statistic improves the efficiency of the more traditional augmented Dickey-Fuller or Phillips-Perron tests, achieving substantially higher power at a cost of only slightly higher size distortions. It is shown to be the most powerful unit root test, especially for small sample sizes and when an unknown mean or trend is present.[37] As it is often argued that the lack of power in small samples, rather than underlying non-stationarity, is the cause of the failure to reject in the Dickey-Fuller type tests, we also consider the Kwiatkowski, Phillips, Schmidt, and Shin (1992) test (KPSS)where the null hypothesis tested is (trend) stationarity.[38]

As Table A.1 shows on the next page, in the levels data, ADF-GLS tests almost uniformly fail to reject the unit root null, while the KPSS tests reject the null of trend stationarity.[39] On the first differenced data, the pattern is the reversed: that non-stationarity IS rejected under ADF-GLS and stationarity cannot be rejected under KPSS. These test results support the view that the series are integrated of order 1.

We then test for the existence of linear cointegration relations among the variables, using the full-system maximum likelihood estimation technique of Johansen (1988) and Johansen and Juselius (1990). This test produces two test statistics. The Trace statistic tests for the at most r cointegrating vectors among a system of $N > r$ time series. The Eigenvalue statistic tests for exactly r cointegrating vectors against the alternative hypothesis of $r + 1$ cointegrating vectors. Although this technique is superior to the Engle-Granger (1987) two-step procedure in detecting cointegration relations, it yields widely dispersed estimates for the cointegrating vectors. Therefore, we choose

[37] See, for instance, Hayashi (2000) and Stock (1994) for the superior performance of this test.

[38] See Engel (2001) for reasons why this argument might not actually be valid. Caner and Kilian (1998) documents large finite-sample size distortions of such tests. Here we conduct the tests as a robustness check for the DF-GLS results.

[39] (The maximum lag order for the test is calculated from the sample size, using the rule provided by Schwert (1989). Approximate 5% critical values are calculated from the response surface estimates of Table 1 in Cheung and Lai (1995), which take both the sample size and the lag specification into account. A note of technicality: If the maximum lag order exceeds one, the optimal lag order is chosen as that which minimizes the Ng-Perron (2001) modified Akaike information criterion (MAIC). Through Monte-Carlo experiments, Ng and Perron established that the MAIC yields substantial size improvements to the DF-GLS test developed in Elliot, Rothenberg and Stock (1996). In addition, they recommend using the MAIC along with GLS detrended data for desirable size and power properties.)

to estimate the cointegrating vector with dynamic OLS.

Table A.2 presents test results for the four commodity price augmented models explained in section 2.1. Some evidence of cointegration is detected in all cases.

B Data Appendix:

Exchange Rates: End of period quarterly nominal exchange rates are taken from IMF's International Financial Statistics (IFS) and Global Financial Database for 1973Q1 to 2001Q2.

Consumer Price Index, Money Stock, and Real Output: Quarterly consumer prices, M1, and GDP volume (1995 = 100) are taken from the IFS. M0 is used for UK.

Short Term Interest Rates: Three-month Treasury bill rates are used as the measure of short-term interest rate. They are taken from Global Financial Database.

Commodity Prices: The country specific commodity export price index is constructed by geometrically weighting the world market prices in US dollar of each country's major non-energy commodity exports. The weights, taken from Djoudad, Murray, Chan, and Daw (2001), are the average production values of the commodities over the 1982-90 period. They are listed in Appendix Table C.1. Note that the following commodities from the original Djoudad *et al* indices are excluded, as we were unable to update the price series. Their original weights in the relevant countries are in the parentheses: barley (2.4% in Australia, 1.8% in Canada), sulphur (1.4% in Canada), cod (0.01% in Canada), lobster (0.5% in Canada), and salmon (0.6% in Canada).

The world market price of individual commodities is taken from various sources (see Table C.1). They are quarterly average spot or cash prices in US dollars. The commodities in general are traded in different markets, including NYMEX, IPE, CBT, CME, KCB, ASX and SFE, and the prices are considered "world prices".

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Table 1. Estimation of the Cointegration Vectors under Dynamic OLS

A. Australian-US Exchange Rate

	PPP		**Asset Approach-Flex Price Monetary**		**Flex Price Monetary**		**Sticky Price Monetary**	
		Comm Price		Comm Price		Comm Price		Comm Price
Commodity Price(+)		*0.92		*0.61		*0.70		*0.56
		(9.49)		(8.26)		(9.61)		(8.34)
Relative CPIs (+)	-0.03	*1.11						
	(-0.10)	(4.21)						
Relative Money Stock (+)			* 0.19	* 0.15	0.22	*0.30	*0.39	*0.36
			(1.83)	(2.07)	(1.59)	(3.95)	(3.76)	(5.24)
Relative Real GDP (-)			-0.95	-0.04	-1.04	-0.91	*-4.57	*-2.56
			(-0.52)	(-0.04)	(-0.52)	(-0.81)	(-2.96)	(-2.50)
Relative Interest Rates					-0.21	*-1.54	*2.25	0.08
(+)/(-)					(-0.31)	(-3.78)	(3.64)	(0.16)
Relative Inflation(+)							*-4.00	*-2.15
							(-7.33)	(-4.92)
Cointegration	Yes	Yes	Yes	Yes	Yes	Yes	No	Yes
Adj. R^2	-0.03	0.61	0.18	0.68	0.13	0.73	0.56	0.84
N. Obs.	68	68	67	67	67	67	67	67

Note:

- A * indicates significance at the 5% level. Heteroskedasticity and autocorrelation consistent t-ratios are reported in the parenthesis. Under the assumption that the variables are cointegrated, DOLS procedure produces efficient estimators for the cointegrating vector. Two leads and lags of each explanatory variable are included in the regressions.
- The theoretical signs for each coefficient are in parentheses following the variables in the table.
- The expected sign on relative interest rates, under the flexible price model and the real interest rate models are positive and negative, respectively.
- Country-specific commodity price index is the base country production-weighted average of world commodity prices in U.S. dollars. See Data Appendix for details.
- "Cointegration" provides results from the Engle-Granger-ADF procedure which tests for the stationarity of the OLS regression residuals.

(Cont.) Table 1. Estimation of the Cointegration Vectors under Dynamic OLS

B. Canadian-US Exchange Rate

	PPP		**Asset Approach-Flex Price Monetary**		**Flex Price Monetary**		**Sticky Price Monetary**	
		Comm Price		Comm Price		Comm Price		Comm Price
Commodity Price(+)		* 0.46		* -0.29		* -0.31		* -0.29
		(8.05)		(-5.72)		(-6.27)		(-5.07)
Relative CPIs (+)	* 1.01	-0.39						
	(3.78)	(-1.51)						
Relative Money Stock (+)			* 0.29	0.07	* 0.33	0.09	* 0.31	** 0.14
			(4.87)	(1.11)	(5.31)	(1.41)	(4.83)	(2.26)
Relative Real GDP (-)			-0.06	-0.51	0.74	0.27	* 0.98	0.25
			(-0.15)	(-1.56)	(1.51)	(0.64)	(1.90)	(0.55)
Relative Interest Rates					* -1.94	* -2.14	* 2.09	* 2.16
(+)/(-)					(-2.43)	(-3.16)	(2.52)	(3.13)
Relative Inflation(+)							* 1.06	-0.34
							(1.67)	(-0.62)
Cointegration	Yes	Yes	No	Yes	No	Yes	No	Yes
Adj. R^2	0.12	0.48	0.54	0.68	0.56	0.70	0.54	0.68
N. Obs.	110	110	108	108	108	108	108	108

Note:

- A * indicates significance at the 5% level. Heteroskedasticity and autocorrelation consistent t-ratios are reported in the parenthesis. Under the assumption that the variables are cointegrated, DOLS procedure produces efficient estimators for the cointegrating vector. Two leads and lags of each explanatory variable are included in the regressions.
- The theoretical signs for each coefficient are in parentheses following the variables in the table.
- The expected sign on relative interest rates, under the flexible price model and the real interest rate models are positive and negative, respectively.
- Country-specific commodity price index is the base country production-weighted average of world commodity prices in U.S. dollars. See Data Appendix for details.
- "Cointegration" provides results from the Engle-Granger-ADF procedure which tests for the stationarity of the OLS regression residuals.

(Cont.) Table 1. Estimation of the Cointegration Vectors under Dynamic OLS

C. New Zealand-US Exchange Rate

	PPP		**Asset Approach-Flex Price Monetary**		**Flex Price Monetary**		**Sticky Price Monetary**	
		Comm Price		Comm Price		Comm Price		Comm Price
Commodity Price(+)		*1.51		*1.44		*1.28		*1.46
		(7.40)		(7.88)		(5.06)		(5.70)
Relative CPIs (+)	*-1.73	-0.49						
	(-3.65)	(-1.30)						
Relative Money Stock (+)			-0.03	*0.29	* 0.27	*0.33	* 0.27	* 0.24
			(-0.24)	(3.10)	(1.95)	(3.02)	(1.86)	(2.25)
Relative Real GDP (-)			*-1.28	-0.19	*-4.69	-1.10	*-5.31	-1.17
			(-2.25)	(-0.48)	(-5.31)	(-1.01)	(-5.55)	(-1.04)
Relative Interest Rates					*4.92	0.99	*-7.27	* -3.85
(+)/(-)					(4.50)	(0.81)	(-3.70)	(-2.55)
Relative Inflation(+)							2.04	*4.03
							(1.01)	(2.80)
Cointegration	No	Yes	No	Yes	No	Yes	No	Yes
Adj. R^2	0.17	0.58	0.20	0.71	0.48	0.73	0.49	0.78
N. Obs.	58	57	57	56	57	56	57	56

Note:
- A * indicates significance at the 5% level. Heteroskedasticity and autocorrelation consistent t-ratios are reported in the parenthesis. Under the assumption that the variables are cointegrated, DOLS procedure produces efficient estimators for the cointegrating vector. Two leads and lags of each explanatory variable are included in the regressions.
- The theoretical signs for each coefficient are in parentheses following the variables in the table.
- The expected sign on relative interest rates, under the flexible price model and the real interest rate models are positive and negative, respectively.
- Country-specific commodity price index is the base country production-weighted average of world commodity prices in U.S. dollars. See Data Appendix for details.
- "Cointegration" provides results from the Engle-Granger-ADF procedure which tests for the stationarity of the OLS regression residuals.

Table 2. Estimation of the Cointegration Vectors under Dynamic OLS

A. Australian-UK Exchange Rate

	PPP		**Asset Approach-Flex Price Monetary**		**Flex Price Monetary**		**Sticky Price Monetary**	
		Comm Price		Comm Price		Comm Price		Comm Price
Commodity Price(+)		* 0.86		* 0.55		0.46		0.27
		(3.06)		(2.43)		(1.63)		(1.00)
Relative CPIs (+)	-0.09	* 1.23						
	(-0.20)	(1.71)						
Relative Money Stock (+)			0.12	* 0.46	0.07	0.33	0.15	0.30
			(1.09)	(2.51)	(0.56)	(1.32)	(1.41)	(1.30)
Relative Real GDP (-)			0.59	-1.70	0.62	-1.15	-0.62	-1.50
			(0.85)	(-1.39)	(0.85)	(-0.80)	(-0.91)	(-1.11)
Relative Interest Rates					1.10	1.23	* 5.69	* 4.97
(-)/(+)					(1.03)	(0.95)	(4.28)	(3.11)
Relative Inflation(+)							*-4.83	*-4.14
							(-4.90)	(-3.43)
Cointegration	No	Yes	Yes	Yes	No	No	No	No
Adj. R^2	-0.10	0.03	0.20	0.28	0.16	0.26	0.40	0.38
N. Obs.	66	66	67	67	67	67	67	67

Note:
- A * indicates significance at the 5% level. Heteroskedasticity and autocorrelation consistent t-ratios are reported in the parenthesis. Under the assumption that the variables are cointegrated, DOLS procedure produces efficient estimators for the cointegrating vector. Two leads and lags of each explanatory variable are included in the regressions.
- The theoretical signs for each coefficient are in parentheses following the variables in the table.
- The expected sign on relative interest rates, under the flexible price model and the real interest rate models are positive and negative, respectively.
- Country-specific commodity price index is the base country production-weighted average of world commodity prices in U.S. dollars. See Data Appendix for details.
- "Cointegration" provides results from the Engle-Granger-ADF procedure which tests for the stationarity of the OLS regression residuals.

(Cont.) Table 2. Estimation of the Cointegration Vectors under Dynamic OLS

B. Canadian-UK Exchange Rate

	PPP		**Asset Approach-Flex Price Monetary**		**Flex Price Monetary**		**Sticky Price Monetary**	
		Comm Price		Comm Price		Comm Price		Comm Price
Commodity Price(+)		-0.01		* -0.23		* -0.25		* -0.29
		(-0.10)		(-2.03)		(-2.23)		(-2.46)
Relative CPIs (+)	-0.02	-0.15						
	(-0.20)	(-0.81)						
Relative Money Stock (+)			0.02	-0.03	-0.02	-0.09	-0.02	* -0.11
			(0.47)	(-0.66)	(-0.36)	(-1.39)	(-0.29)	(-1.67)
Relative Real GDP (-)			0.31	-0.20	* 0.89	-0.06	0.31	-0.27
			(1.12)	(-0.48)	(2.12)	(-0.13)	(0.83)	(-0.53)
Relative Interest Rates (+)/(-)					0.50	1.32	1.16	* 1.91
					(0.43)	(1.38)	(1.08)	(1.83)
Relative Inflation(+)							-0.58	-0.70
							(-1.29)	(-1.58)
Cointegration	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adj. R^2	-0.03	0.09	0.06	0.13	0.11	0.16	0.05	0.15
N. Obs.	112	112	110	110	110	110	110	110

Note:

- A * indicates significance at the 5% level. Heteroskedasticity and autocorrelation consistent t-ratios are reported in the parenthesis. Under the assumption that the variables are cointegrated, DOLS procedure produces efficient estimators for the cointegrating vector. Two leads and lags of each explanatory variable are included in the regressions.
- The theoretical signs for each coefficient are in parentheses following the variables in the table.
- The expected sign on relative interest rates, under the flexible price model and the real interest rate models are positive and negative, respectively.
- Country-specific commodity price index is the base country production-weighted average of world commodity prices in U.S. dollars. See Data Appendix for details.
- "Cointegration" provides results from the Engle-Granger-ADF procedure which tests for the stationarity of the OLS regression residuals.

(Cont.) Table 2. Estimation of the Cointegration Vectors under Dynamic OLS

C. New Zealand-UK Exchange Rate

	PPP		**Asset Approach-Flex Price Monetary**		**Flex Price Monetary**		**Sticky Price Monetary**	
		Comm Price		Comm Price		Comm Price		Comm Price
Commodity Price(+)		* 1.92		* 1.61		* 1.32		* 1.28
		(9.92)		(8.42)		(4.69)		(4.43)
Relative CPIs (+)	-0.32	* 0.47						
	(-0.99)	(2.25)						
Relative Money Stock (+)			-0.05	* 0.32	*-0.65	-0.03	*-0.58	-0.06
			(-0.34)	(2.57)	(-3.80)	(-0.15)	(-2.55)	(-0.22)
Relative Real GDP (-)			-0.56	0.14	0.50	0.46	0.47	0.37
			(-0.87)	(0.30)	(0.90)	(0.95)	(0.85)	(0.74)
Relative Interest Rates (+)/(-)					* 4.75	* 1.83	* 5.54	* 2.04
					(5.68)	(1.78)	(5.27)	(1.73)
Relative Inflation(+)							-0.76	0.01
							(-0.78)	(0.01)
Cointegration	No	Yes	No	Yes	No	Yes	No	Yes
Adj. R^2	-0.02	0.65	-0.10	0.54	0.31	0.55	0.32	0.53
N. Obs.	58	57	59	58	59	58	59	58

Note:

- A * indicates significance at the 5% level. Heteroskedasticity and autocorrelation consistent t-ratios are reported in the parenthesis. Under the assumption that the variables are cointegrated, DOLS procedure produces efficient estimators for the cointegrating vector. Two leads and lags of each explanatory variable are included in the regressions.
- The theoretical signs for each coefficient are in parentheses following the variables in the table.
- The expected sign on relative interest rates, under the flexible price model and the real interest rate models are positive and negative, respectively.
- Country-specific commodity price index is the base country production-weighted average of world commodity prices in U.S. dollars. See Data Appendix for details.
- "Cointegration" provides results from the Engle-Granger-ADF procedure which tests for the stationarity of the OLS regression residuals.

Table 3. Estimation of the Cointegration Vectors under Dynamic OLS

A. Australian-Japan Exchange Rate

	PPP		**Asset Approach-Flex Price Monetary**		**Flex Price Monetary**		**Sticky Price Monetary**	
		Comm Price		Comm Price		Comm Price		Comm Price
Commodity Price(+)		* 0.46		* 0.39		* 0.71		0.18
		(3.88)		(2.09)		(2.44)		(1.03)
Relative CPIs (+)	* 2.03	* 2.02						
	(14.89)	(19.85)						
Relative Money Stock (+)			* 0.86	* 1.04	* 0.88	* 1.28	* 1.04	* 1.08
			(8.68)	(13.58)	(5.67)	(6.82)	(13.85)	(10.96)
Relative Real GDP (-)			0.42	-0.12	0.35	-0.77	*-0.78	*-0.73
			(1.06)	(-0.43)	(0.69)	(-1.49)	(-3.03)	(-2.35)
Relative Interest Rates (+)/(-)					-0.31	-2.59	* 3.72	* 2.15
					(-0.19)	(-1.67)	(4.64)	(2.20)
Relative Inflation(+)							*-7.66	*-5.27
							(-10.95)	(-6.31)
Cointegration	Yes	Yes	Yes	Yes	No	Yes	Yes	Yes
Adj. R^2	0.77	0.88	0.63	0.86	0.58	0.85	0.88	0.91
N. Obs.	68	68	65	65	65	65	67	67

Note:
- A * indicates significance at the 5% level. Heteroskedasticity and autocorrelation consistent t-ratios are reported in the parenthesis. Under the assumption that the variables are cointegrated, DOLS procedure produces efficient estimators for the cointegrating vector. Two leads and lags of each explanatory variable are included in the regressions.
- The theoretical signs for each coefficient are in parentheses following the variables in the table.
- The expected sign on relative interest rates, under the flexible price model and the real interest rate models are positive and negative, respectively.
- Country-specific commodity price index is the base country production-weighted average of world commodity prices in U.S. dollars. See Data Appendix for details.
- "Cointegration" provides results from the Engle-Granger-ADF procedure which tests for the stationarity of the OLS regression residuals.

(Cont.) Table 3. Estimation of the Cointegration Vectors under Dynamic OLS

B. Canadian-Japan Exchange Rate

	PPP		**Asset Approach-Flex Price Monetary**		**Flex Price Monetary**		**Sticky Price Monetary**	
		Comm Price		Comm Price		Comm Price		Comm Price
Commodity Price(+)		-0.25		* -1.01				* -0.76
		(-1.54)		(-7.29)				(-3.59)
Relative CPIs (+)	* 2.10	* 1.98						
	(25.84)	(12.86)						
Relative Money Stock (+)			* 1.31	* 0.99	* 1.44	* 0.88	* 1.42	* 1.04
			(11.53)	(10.19)	(13.08)	(5.75)	(14.08)	(8.28)
Relative Real GDP (-)			* -1.31	-0.27	* -1.16	-0.31	* -1.08	* -0.62
			(-4.05)	(-0.82)	(-3.70)	(-0.77)	(-3.77)	(-1.79)
Relative Interest Rates (+)/(-)					* -5.58	-0.98	-1.58	1.69
					(-5.49)	(-0.69)	(-1.33)	(1.22)
Relative Inflation(+)						* -1.12	* -2.64	* -2.88
						(-4.13)	(-3.31)	(-3.85)
Cointegration	Yes	Yes	Yes	Yes	No	Yes	No	Yes
Adj. R^2	0.86	0.89	0.80	0.88	0.85	0.89	0.86	0.89
N. Obs.	112	112	110	110	108	108	110	110

Note:
- A * indicates significance at the 5% level. Heteroskedasticity and autocorrelation consistent t-ratios are reported in the parenthesis. Under the assumption that the variables are cointegrated, DOLS procedure produces efficient estimators for the cointegrating vector. Two leads and lags of each explanatory variable are included in the regressions.
- The theoretical signs for each coefficient are in parentheses following the variables in the table.
- The expected sign on relative interest rates, under the flexible price model and the real interest rate models are positive and negative, respectively.
- Country-specific commodity price index is the base country production-weighted average of world commodity prices in U.S. dollars. See Data Appendix for details.
- "Cointegration" provides results from the Engle-Granger-ADF procedure which tests for the stationarity of the OLS regression residuals.

(Cont.) Table 3. Estimation of the Cointegration Vectors under Dynamic OLS

C. New Zealand-Japan Exchange Rate

	PPP		**Asset Approach-Flex Price Monetary**		**Flex Price Monetary**		**Sticky Price Monetary**	
		Comm Price		Comm Price		Comm Price		Comm Price
Commodity Price(+)		* 1.10 (4.08)		* 0.77 (2.55)		-0.46 (-1.55)		* -0.93 (-2.50)
Relative CPIs (+)	* 1.99 (4.78)	* 2.10 (6.40)						
Relative Money Stock (+)			* 0.39 (2.21)	* 0.40 (2.24)	0.03 (0.20)	-0.28 (-1.59)	0.03 (0.24)	* -0.37 (-2.01)
Relative Real GDP (-)			* 0.81 (2.19)	* 0.85 (2.36)	* 1.74 (5.64)	* 1.74 (5.74)	* 1.76 (5.16)	* 1.76 (5.51)
Relative Interest Rates (+)/(-)					* 4.73 (6.62)	* 5.80 (6.21)	* 5.38 (5.57)	* 7.97 (5.48)
Relative Inflation(+)							-0.65 (-0.99)	* -1.68 (-2.08)
Cointegration	Yes	Yes	Yes	Yes	Yes	Yes	No	Yes
Adj. R^2	0.35	0.63	0.42	0.69	0.72	0.75	0.70	0.75
N. Obs.	58	57	59	58	59	58	59	58

Note:
- A * indicates significance at the 5% level. Heteroskedasticity and autocorrelation consistent t-ratios are reported in the parenthesis. Under the assumption that the variables are cointegrated, DOLS procedure produces efficient estimators for the cointegrating vector. Two leads and lags of each explanatory variable are included in the regressions.
- The theoretical signs for each coefficient are in parentheses following the variables in the table.
- The expected sign on relative interest rates, under the flexible price model and the real interest rate models are positive and negative, respectively.
- Country-specific commodity price index is the base country production-weighted average of world commodity prices in U.S. dollars. See Data Appendix for details.
- "Cointegration" provides results from the Engle-Granger-ADF procedure which tests for the stationarity of the OLS regression residuals.

Table 4. Australian-US In-Sample Predictive Regressions: Slope Coefficient λ_k and Associated t-Statistics

$\Delta s_{t+k} = \alpha_k + \lambda_k(\beta_k f_t - s_t) + \varepsilon_{t+k}$

Two alternative constraints on the cointegrating vector β

	K = 1							
Constraints:	PPP	PPP + P^{Com}	MON1	MON1+ P^{Com}	MON2	MON2+ P^{Com}	MON3	MON3+ P^{Com}
1) $\beta_{cp} = 0.8$; $\beta_p = 1$; $\beta_m = 1$ $\beta_y = -1$; $\beta_i = +/-1$; $\beta_\pi = 1$	0.02 (0.52)	* 0.18 (2.34)	0.03 (1.34)	* 0.04 (1.83)	0.03 (1.41)	* 0.04 (1.82)	0.03 (1.45)	* 0.04 (1.87)
Adj R^2	-0.01	0.06	0.01	0.03	0.01	0.03	0.02	0.04
2) β = Estimated Cointegrating Vectors	0.08 (1.52)	* 0.19 (2.40)	* 0.14 (2.28)	* 0.31 (4.20)	* 0.14 (2.27)	* 0.32 (4.11)	0.05 (1.27)	* 0.22 (2.98)
Adj R^2	0.02	0.07	0.06	0.20	0.06	0.19	0.01	0.10
	K = 4 (w/ Newey-West t-stats)							
1) $\beta_{cp} = 0.8$; $\beta_m = 1$ $\beta_y = -1$; $\beta_i = +/-1$; $\beta_\pi = 1$	0.15 (0.97)	* 0.77 (2.87)	* 0.12 (2.01)	* 0.14 (2.46)	*1.12 (2.12)	* 1.13 (2.53)	* 0.13 (2.18)	* 0.16 (2.71)
Adj R^2	0.01	0.22	0.11	0.15	0.11	0.15	0.12	0.17
2) β = Estimated Cointegrating Vectors	0.40 (1.63)	* 0.71 (3.11)	* 0.59 (3.21)	*1.19 (11.97)	* 0.58 (3.26)	* 1.11 (7.18)	0.13 (1.27)	* 0.64 (4.46)
Adj R^2	0.12	0.21	0.25	0.62	0.24	0.51	0.01	0.20
N. Obs	69, 66 (for K = 1, 4)							

Note: A * indicates significance at the 5% level. Heteroskedasticity and autocorrelation consistent t-ratios are reported in the parenthesis. Regressions based on the models below are conducted with and without the commodity price term. The "Estimated Cointegrating Vectors" for β are taken from Table 1-3.

- PPP: $\Delta s_{t+k} = \alpha + \lambda_k [\beta_{cp} p_t^{com} + \beta_p(p_t^* - p_t) - s_t] + \varepsilon_t$
- MON1: $\Delta s_{t+k} = \alpha + \lambda_k [\beta_{cp} p_t^{com} + \beta_m(m_t^* - m_t) + \beta_y(y_t^* - y_t) - s_t] + \varepsilon_t$
- MON2: $\Delta s_{t+k} = \alpha + \lambda_k [\beta_{cp} p_t^{com} + \beta_m(m_t^* - m_t) + \beta_y(y_t^* - y_t) + \beta_i(i_t^* - i_t) - s_t] + \varepsilon_t$
- MON3: $\Delta s_{t+k} = \alpha + \lambda_k [\beta_{cp} p_t^{com} + \beta_m(m_t^* - m_t) + \beta_y(y_t^* - y_t) + \beta_i(i_t^* - i_t) + \beta_\pi(\pi_t^* - \pi_t) - s_t] + \varepsilon_t$

(Cont.) Canada-US In-Sample Predictive Regressions: Slope Coefficient λ_k and Associated t-Statistics

$\Delta s_{t+k} = \alpha_k + \lambda_k(\beta_k f_t - s_t) + \varepsilon_{t+k}$; with two alternative constraints on β

	K = 1							
Constraints:	PPP	PPP + P^{Com}	MON1	MON1+ P^{Com}	MON2	MON2+ P^{Com}	MON3	MON3+ P^{Com}
1) β_{cp} = 0.3; β_m = 1	0.01	0.01	-0.00	0.00	0.00	0.00	-0.00	-0.00
β_y = -1; β_i = +/-1; β_π = 1	(0.78)	(0.81)	(-0.12)	(0.10)	(0.04)	(0.26)	(-0.38)	(-0.19)
Adj R^2	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00
2) β = Estimated	0.01	0.01	0.02	0.03	-0.00	-0.00	0.02	* 0.05
Cointegrating Vectors	(0.78)	(0.92)	(0.89)	(1.06)	(-0.01)	(-0.00)	(1.08)	(1.78)
Adj R^2	-0.00	-0.00	-0.00	0.00	-0.00	-0.00	0.00	0.02
	K = 4 (w/ Newey-West t-stats)							
1) β_{cp} = 0.3; β_m = 1	0.07	0.04	0.00	0.01	0.00	0.01	0.00	0.01
β_y = -1; β_i = +/-1; β_π = 1	(1.04)	(0.86)	(0.04)	(0.20)	(0.12)	(0.28)	(0.05)	(0.21)
Adj R^2	0.02	0.01	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00
2) β = Estimated	0.07	0.03	0.10	0.15	0.03	0.05	0.09	* 0.15
Cointegrating Vectors	(1.04)	(0.90)	(1.16)	(1.63)	(0.49)	(0.74)	(1.27)	(1.69)
Adj R^2	0.02	0.01	0.02	0.06	-0.00	0.01	0.03	0.05
	K = 8 (w/ Newey-West t-stats)							
1) β_{cp} = 0.3; β_m = 1	* 0.21	* 0.12	0.02	0.03	0.02	0.03	0.02	0.03
β_y = -1; β_i = +/-1; β_π = 1	(1.94)	(0.68)	(0.32)	(0.59)	(0.35)	(0.64)	(0.30)	(0.55)
Adj R^2	0.09	0.06	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00
2) β = Estimated	*0.21	0.08	* 0.27	* 0.40	0.15	0.18	* 0.22	* 0.38
Cointegrating Vectors	(1.94)	(1.57)	(1.94)	(2.43)	(1.33)	(1.51)	(1.75)	(2.31)
Adj R^2	0.09	0.06	0.12	0.21	0.05	0.08	0.08	0.17
N. Obs	113, 110, 106 (For K = 1, 4, 8)							

(Cont.) New Zealand-US In-Sample Predictive Regressions: Slope Coefficient λ_k and Associated t-Statistics

$\Delta s_{t+k} = \alpha_k + \lambda_k(\beta_k f_t - s_t) + \varepsilon_{t+k}$

with two alternative constraints on β

	K = 1							
Constraints:	PPP	PPP + P^{Com}	MON1	MON1+ P^{Com}	MON2	MON2+ P^{Com}	MON3	MON3+ P^{Com}
1) β_{cp} = 1.4; β_m = 1	0.00	0.00	* 0.05	* 0.07	* 0.05	* 0.06	* 0.05	* 0.06
β_y = -1; β_i = +/-1; β_π = 1	(0.05)	(0.04)	(2.16)	(2.42)	(2.31)	(2.51)	(2.17)	(2.42)
Adj R^2	-0.00	-0.00	0.06	0.07	0.07	0.08	0.06	0.07
2) β = Estimated	0.04	0.04	0.04	0.10	* 0.03	* 0.09	0.01	0.04
Cointegrating Vectors	(0.77)	(0.48)	(0.82)	(1.63)	(2.15)	(2.24)	(0.36)	(1.47)
Adj R^2	-0.00	-0.00	-0.00	0.03	0.06	0.06	-0.00	0.02
	K = 4 (w/ Newey-West t-stats)							
1) β_{cp} = 1.4; β_m = 1	0.22	* 0.48	* 0.29	* 0.35	* 0.26	* 0.30	* 0.27	* 0.33
β_y = -1; β_i = +/-1; β_π = 1	(0.91)	(1.65)	(4.89)	(6.75)	(5.01)	(6.47)	(5.40)	(7.08)
Adj R^2	0.04	0.09	0.35	0.42	0.35	0.40	0.35	0.42
2) β = Estimated	0.24	* 0.38	* 0.45	* 0.85	* 0.14	* 0.50	* 0.14	* 0.23
Cointegrating Vectors	(1.00)	(1.79)	(1.70)	(4.81)	(3.59)	(6.12)	(1.75)	(6.38)
Adj R^2	0.03	0.05	0.16	0.40	0.24	0.39	0.08	0.20
N. Obs	61, 58 (for K = 1, 4)							

Note: A * indicates significance at the 5% level. Heteroskedasticity and autocorrelation consistent t-ratios are reported in the parenthesis. Regressions based on the models below are conducted with and without the commodity price term. The "Estimated Cointegrating Vectors" for β are taken from Table 1-3.

- PPP: $\Delta s_{t+k} = \alpha + \lambda_k [\beta_{cp} p_t^{com} + \beta_p(p_t^* - p_t) - s_t] + \varepsilon_t$
- MON1: $\Delta s_{t+k} = \alpha + \lambda_k [\beta_{cp} p_t^{com} + \beta_m(m_t^* - m_t) + \beta_y(y_t^* - y_t) - s_t] + \varepsilon_t$
- MON2: $\Delta s_{t+k} = \alpha + \lambda_k [\beta_{cp} p_t^{com} + \beta_m(m_t^* - m_t) + \beta_y(y_t^* - y_t) + \beta_i(i_t^* - i_t) - s_t] + \varepsilon_t$
- MON3: $\Delta s_{t+k} = \alpha + \lambda_k [\beta_{cp} p_t^{com} + \beta_m(m_t^* - m_t) + \beta_y(y_t^* - y_t) + \beta_i(i_t^* - i_t) + \beta_\pi(\pi_t^* - \pi_t) - s_t] + \varepsilon_t$

Table 5. Representative Exchange Rate Predictability Results using Kilian (1999) Bootstrap

"Constrained models" with Theory-Based Cointegrating Vectors Imposed

$$\Delta s_{t+k} = \alpha + \lambda[\beta f_t - s_t] + \varepsilon_{t+k}$$

AUS-Japan Exchange Rate		t-stats for $\lambda > 0$	RMSE Ratio (Model/RW)	Diebold-Mariano Statistics
Fundamental =	k = 1	4.82	0.95	2.68
Relative CPI +		(p = 0.04)	(p = 0.02)	(p = 0.02)
0.8 * Commodity Price	k = 4	9.34	0.82	1.99
		(p = 0.01)	(p = 0.02)	(p = 0.06)
	k = 8	20.16	0.65	2.03
		(p = 0.00)	(p = 0.02)	(p = 0.04)

AUS-Japan Exchange Rate				
Fundamental =	k = 1	3.38	0.93	1.43
Relative Money -		(p = 0.02)	(p = 0.01)	(p = 0.02)
Relative Output +	k = 4	5.48	0.77	1.74
0.8* Commodity Price		(p = 0.02)	(p = 0.01)	(p = 0.03)
	k = 8	7.28	0.85	0.72
		(p = 0.02)	(p = 0.09)	(p = 0.14)

NZL-US Exchange Rate				
Fundamental =	k = 1	4.37	0.97	1.47
Relative Money -		(p = 0.03)	(p = 0.04)	(p = 0.01)
Relative Output +	k = 4	5.68	0.86	1.79
1.4* Commodity Price		(p = 0.02)	(p = 0.04)	(p = 0.03)
	k = 8	4.28	0.95	0.74
		(p = 0.08)	(p = 0.20)	(p = 0.15)

NZL-US Exchange Rate				
Fundamental =	k = 1	4.72	0.97	1.61
Relative Money -		(p = 0.02)	(p = 0.05)	(p = 0.02)
Relative Output -	k = 4	7.85	0.87	1.84
Relative Interest Rates +		(p = 0.00)	(p = 0.05)	(p = 0.03)
Relative Inflation +	k = 8	4.97	0.96	0.70
1.4* Commodity Price		(p = 0.03)	(p = 0.20)	(p = 0.16)

NZL-Japan Exchange Rate				
Fundamental =	k = 1	3.66	0.92	3.70
Relative CPI +		(p = 0.02)	(p = 0.01)	(p = 0.01)
1.4* Commodity Price	k = 4	6.48	0.66	7.56
		(p = 0.02)	(p = 0.00)	(p = 0.00)
	k = 8	8.74	0.51	10.39
		(p = 0.04)	(p = 0.01)	(p = 0.00)

Appendix Table A1. Representative Unit Root Test Statistics

A. In Levels

	Australia		*Canada*		*New Zealand*	
	DF-GLS	KPSS	DF-GLS	KPSS	DF-GLS	KPSS
Log(Nominal ExRate)	-2.13 [-3.01]	0.17** [0.15]	-1.75 [-3.01]	0.15** [0.15]	-1.93 [-3.01]	0.25*** [0.15]
Log(CPI)	-0.83 [-2.94]	0.37*** [0.15]	-1.01 [-2.96]	0.38*** [0.15]	-0.26 [-3.00]	0.38*** [0.15]
Log(Non-Energy Commodity Price)	-2.42 [3.08]	0.14* [0.15]	-0.77 [-2.83]	0.24*** [0.15]	-2.38 [-3.12]	0.21** [0.15]
Log(Energy Commodity Price)	-1.81 [-2.81]	0.09 [0.15]	-0.75 [-3.00]	0.28*** [0.15]	NA	NA
Log (M1)	-1.78 [-3.02]	0.22*** [0.15]	-1.66 [-3.00]	0.12* [0.15]	-1.28 [-3.02]	0.17** [0.15]
Interest Rate	-1.78 [-2.96]	0.30*** [0.15]	-1.34 [-2.96]	0.26*** [0.15]	-2.25 [-3.01]	0.19** [0.15]
Log(real GDP)	-1.62 [-3.02]	0.18** [0.15]	-1.87 [-3.02]	0.15** [0.15]	-1.52 [-3.08]	0.20** [0.15]

B. In First Differences

	Australia		*Canada*		*New Zealand*	
	DF-GLS	KPSS	DF-GLS	KPSS	DF-GLS	KPSS
ΔLog(Nominal ExRate)	-6.25*** [-3.01]	0.08 [0.15]	-4.14*** [-2.99]	0.07 [0.15]	-4.71*** [-2.99]	0.07 [0.15]
ΔLog(CPI]	-3.46** [-2.96]	0.07 [0.15]	-2.71* [-2.98]	0.09 [0.15]	-3.87*** [-3.02]	0.11 [0.15]
ΔLog(Non-Energy Commodity Price)	-3.16** [-3.11]	0.05 [0.15]	-3.77*** [-2.99]	0.06 [0.15]	-2.44 [-3.17]	0.11 [0.15]
ΔLog(Energy Commodity Price]	-2.05 [-3.00]	0.07 [0.15]	-4.46*** [-2.96]	0.15** [0.15]	NA	NA
ΔLog (M1)	-4.18*** [-3.02]	0.05 [0.15]	-3.26** [-2.96]	0.07 [0.15]	-3.17* [3.00]	0.07 [0.15]
Δ(Interest Rate)	-3.90*** [-2.98]	0.05 [0.15]	-5.15*** [-2.99]	0.06 [0.15]	-5.61*** [-3.05]	0.08 [0.15]
ΔLog(real GDP)	-4.22*** [-3.02]	0.05 [0.15]	-5.00*** [-3.02]	0.08 [0.15]	-2.75 [-3.08]	0.11 [0.15]

Notes:
1. The 5% critical values for the optimal lag length tested are reported in the square brackets. *** indicates rejection of the null at 1% significance level; ** at 5% and * at 10%. Maximum lag length tested is chosen based on the Schwert criterion (Schwert 1989). Statistics for Ng-Perron (2001) MAIC lag length is reported.
2. Exchange Rates are measured relative to the U.S. dollar.
3. For the DF-GLS, the null being tested is non-stationarity (with a trend?), and the critical values are based on Cheung and Lai (1995) which accounts for both the sample size and the lag specification.
4. For the KPSS test, the null hypothesis being tested is trend stationarity.
5. Sample sizes vary depending on data availability; most series start in 1973Q1, giving roughly a hundred data points. The shortest series is the commodity price series for New Zealand, with just 50 data points.

Fig. 1a: US-Australian Nominal Exchange Rate, Relative CPI, and Commodity Price Index, 1983q1 to 2001q2

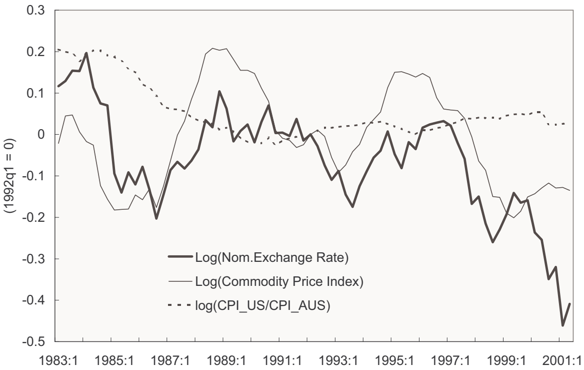


Fig. 1b: Japan-Australian Nominal Exchange Rate, Relative CPI, and Commodity Price Index, 1983q1 to 2001q2

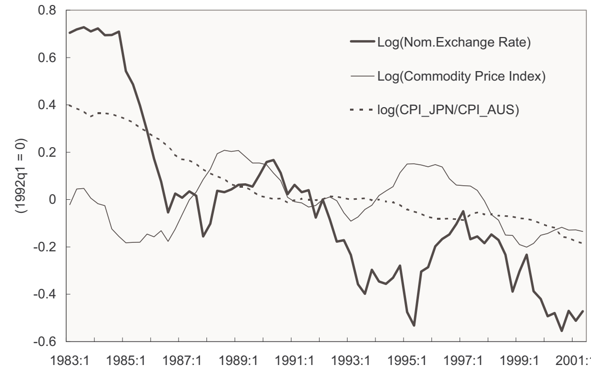
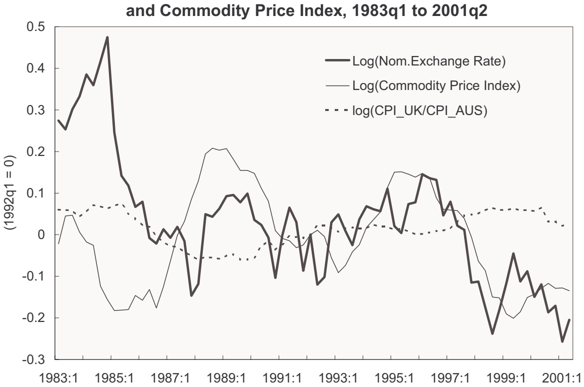


Fig. 1c: UK - Australian Nominal Exchange Rate, Relative CPI, and Commodity Price Index, 1983q1 to 2001q2



Notes:
1. Nominal exchange rates are defined as the foreign currency price for a unit of home currency. A higher number thus represents an appreciation of the home currency.

2. The commodity price index is the averaged world price in US dollars of the country's main commodity exports, using home production shares as weights. See Data Appendix for details.

3. Relative CPIs are log CPI in the anchor country relative to log CPI in the home country.

Fig. 2a: US - Canadian Nominal Exchange Rate, Relative CPI, and Commodity Price Index, 1973q1 to 2001q2

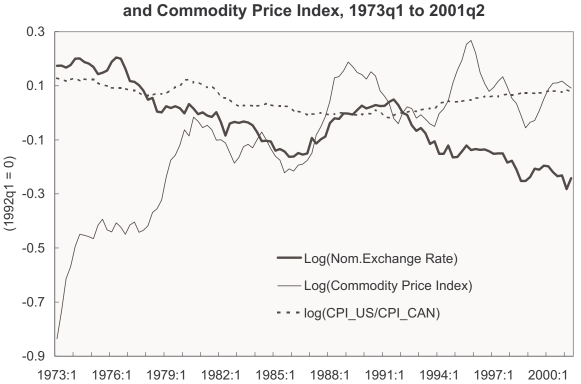


Fig. 2b: Japan - Canadian Nominal Exchange Rate, Relative CPI, and Commodity Price Index, 1973q1 to 2001q2

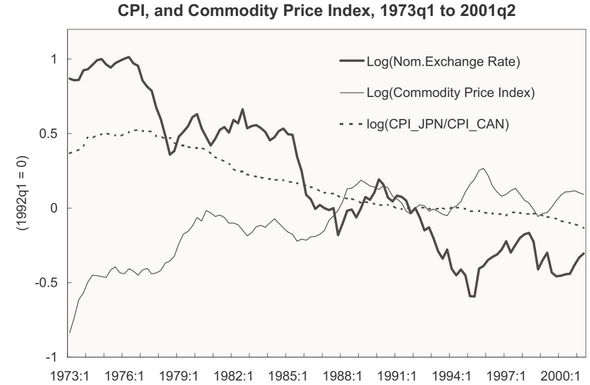
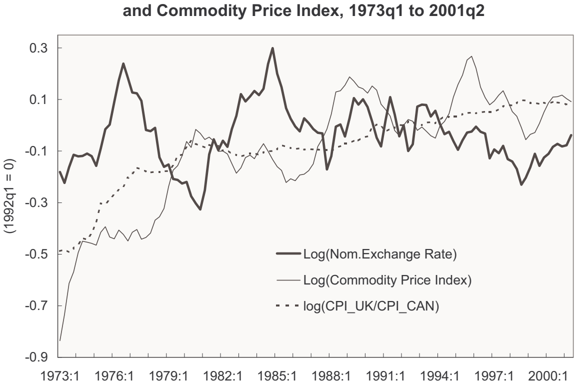


Fig. 2c: UK - Canadian Nominal Exchange Rate, Relative CPI, and Commodity Price Index, 1973q1 to 2001q2



Notes:
1. Nominal exchange rates are defined as the foreign currency price for a unit of home currency. A higher number thus represents an appreciation of the home currency.

2. The commodity price index is the averaged world price in US dollars of the country's main commodity exports, using home production shares as weights. See Data Appendix for details.

3. Relative CPIs are log CPI in the anchor country relative to log CPI in the home country.

Fig. 3a: US - New Zealand Nominal Exchange Rate, Relative CPI, and Commodity Price Index, 1986q1 to 2001q2

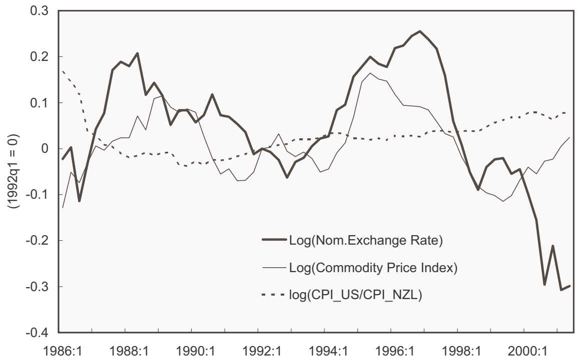


Fig. 3b: Japan - New Zealand Nominal Exchange Rate, Relative CPI, and Commodity Price Index, 1986q1 to 2001q2

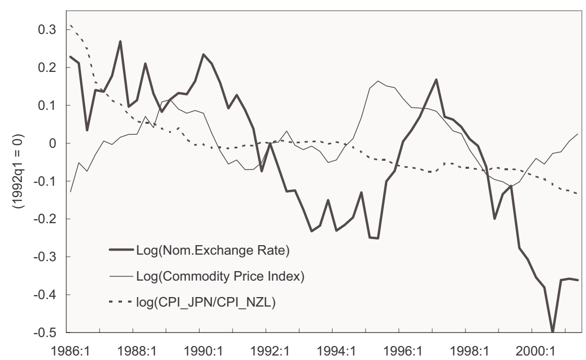
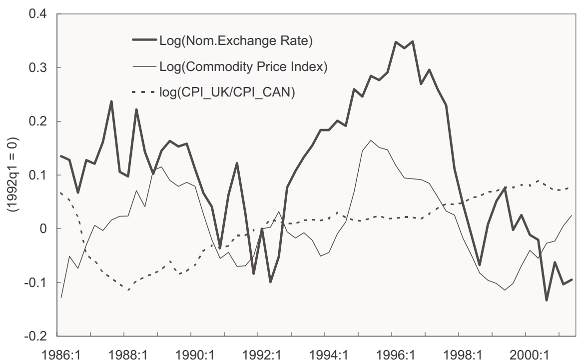


Fig. 3c: UK - New Zealand Nominal Exchange Rate, Relative CPI, and Commodity Price Index, 1986q1 to 2001q2



Notes:

1. Nominal exchange rates are defined as the foreign currency price for a unit of home currency. A higher number thus represents an appreciation of the home currency.

2. The commodity price index is the averaged world price in US dollars of the country's main commodity exports, using home production shares as weights. See Data Appendix for details.

3. Relative CPIs are log CPI in the anchor country relative to log CPI in the home country.

Figure 4. Simulated Out-of-Sample Forecast Comparisons

RMSE Ratios: Models vs. Random Walk

Constrained Specifications with and without Commodity Prices

(See Charts on Next Three Pages)

Note:

1. The charts present ratios of root mean square forecast errors (RMSEs) produced by the model relative to those produced by the random walk. The models (PPP, Mon1, Mon2, & Mon3) are estimated **with cointegrating parameter constraints pre-imposed**, according to the specifications below:

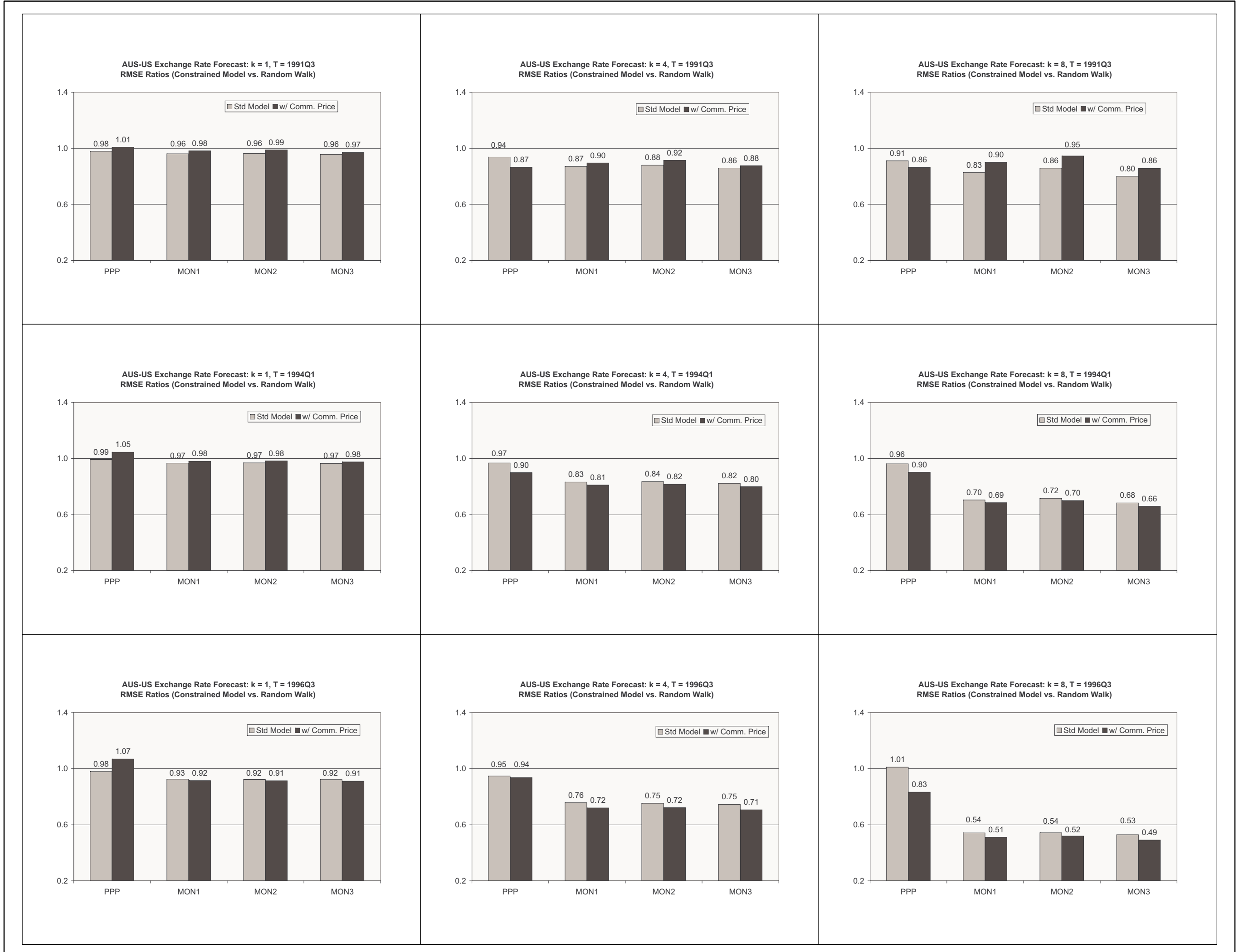
2. The forecast equations are (with and without the $\beta_{cp}p_t^{com}$ term):
   - PPP: $\Delta s_{t+k} = \alpha + \lambda_k[(p_t^* / p_t) + \beta_{cp}p_t^{com} - s_t] + \varepsilon_{t+k}$
   - MON1: $\Delta s_{t+k} = \alpha + \lambda_k [(m_t^* / m_t) - (y_t^* / y_t) + \beta_{cp}p_t^{com} - s_t] + \varepsilon_{t+k}$
   - MON2: $\Delta s_{t+k} = \alpha + \lambda_k [(m_t^*/ m_t) - (y_t^* / y_t) + (i_t^* - i_t) + \beta_{cp}p_t^{com} - s_t] + \varepsilon_{t+k}$
   - MON3: $\Delta s_{t+k} = \alpha + \lambda_k [(m_t^* / m_t) - (y_t^* / y_t) - (i_t^* - i_t) + (\pi_t^* - \pi_t) + \beta_{cp}p_t^{com} - s_t] + \varepsilon_{t+k}$

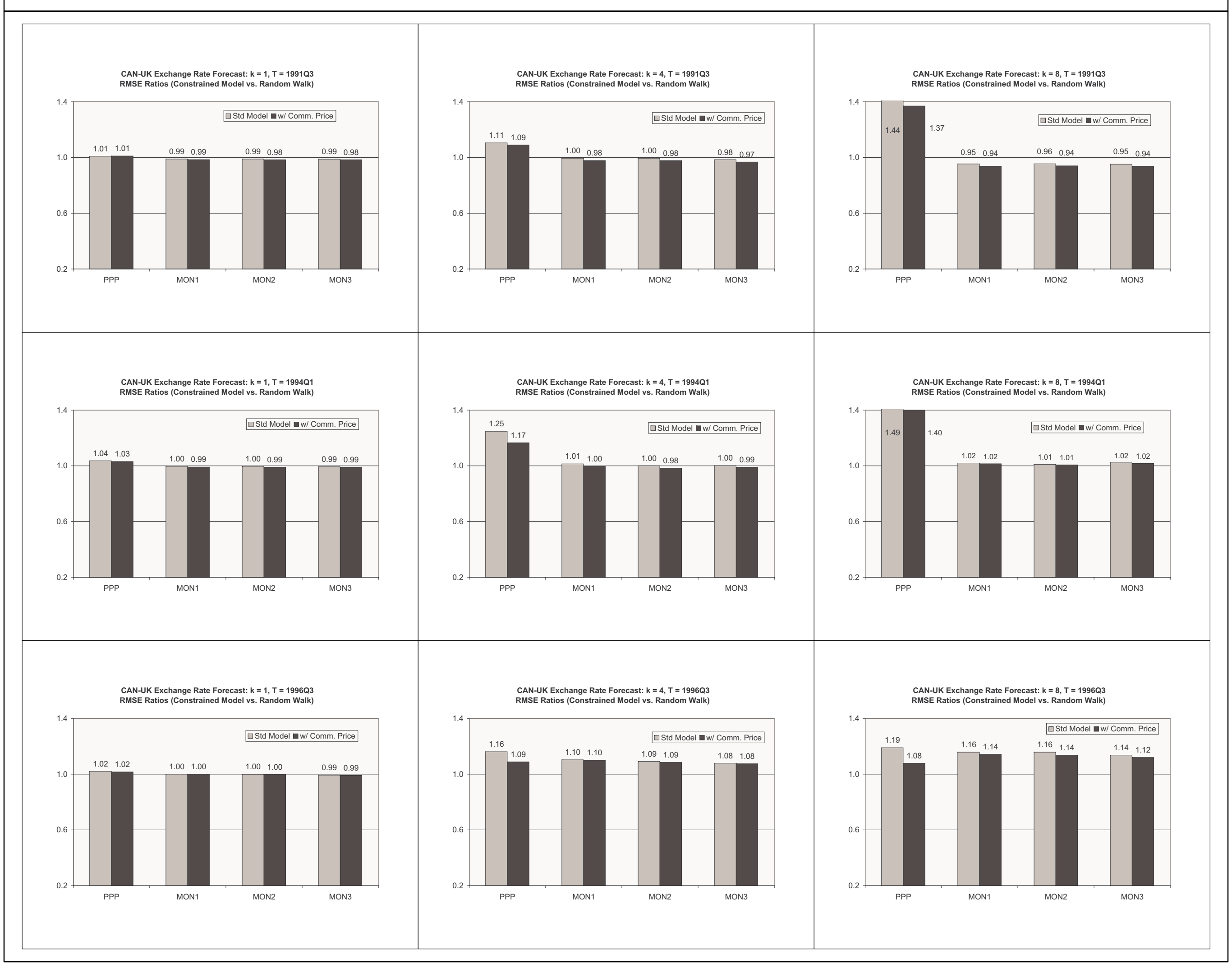
   where β_{cp} is 0.8, -0.3, and 1.4 for Australia, Canada, and New Zealand, respectively. k is the forecast horizon. All variables are in logs except for short-term interest rates and inflation rates.

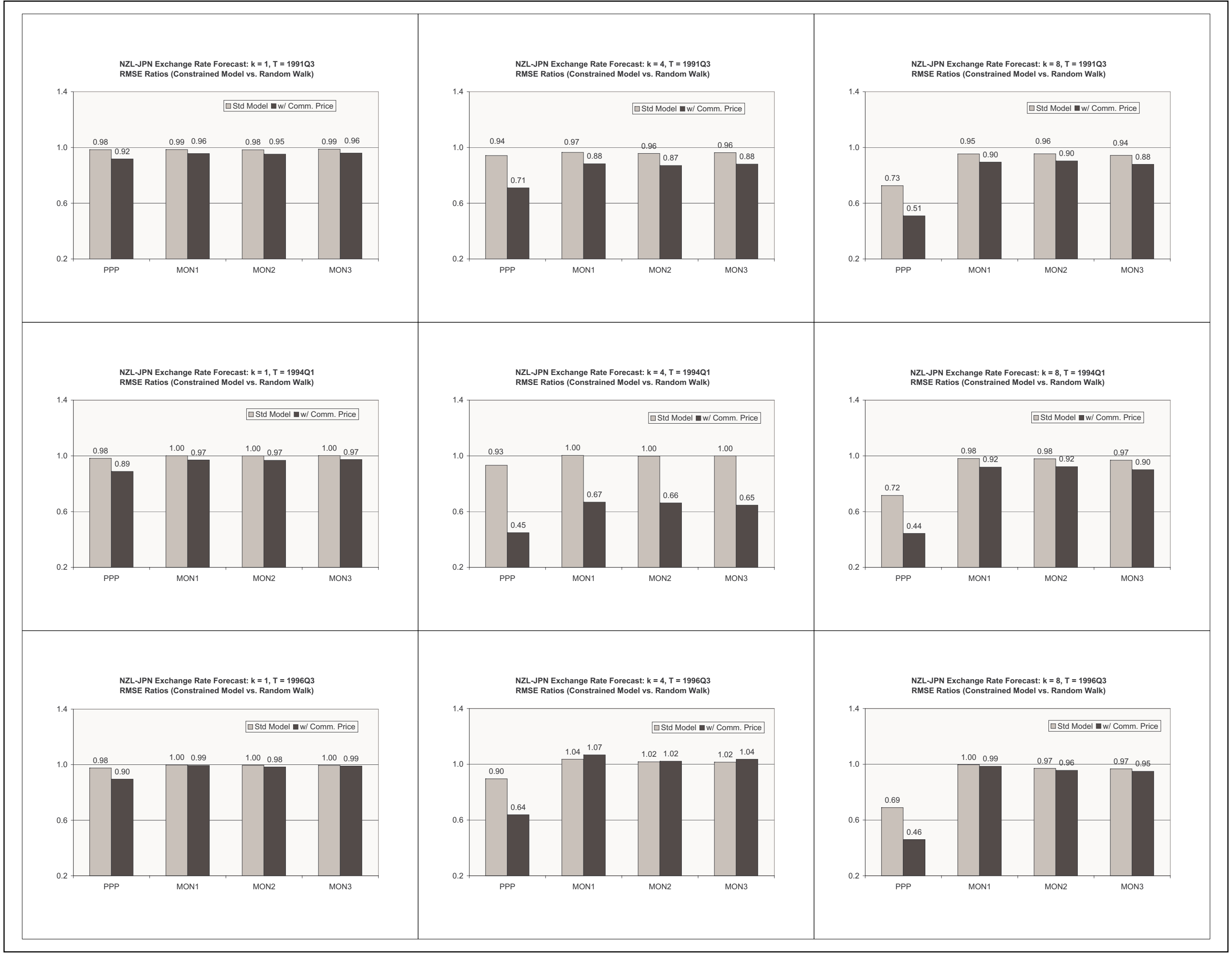
3. The lighter bars are RMSE ratios (relative to RW forecast) of forecast results from "Standard Models" without including commodity prices in the above equations; the darker bars are from models including commodity prices as specified above.

4. The nine charts on each page are organized as follows:
   a. From left to right: forecast horizon k increases from 1, 4, to 8
   b. From top to bottom: forecast period decreases from the last 40 to 30 to 20 observations. (This corresponds to forecast starting points: 1991Q3, 1994Q1, and 1996Q3.)

5. See texts for more discussions. Three representative currency pairs are presented here only; results for other 6 currency pairs are available upon request.


AUS-US Exchange Rate Forecast: k = 1, T = 1991Q3
RMSE Ratios (Constrained Model vs. Random Walk)
Std Model
w/ Comm. Price
0.98 1.01
0.96 0.98
0.96 0.99
0.96 0.97
AUS-US Exchange Rate Forecast: k = 4, T = 1991Q3
RMSE Ratios (Constrained Model vs. Random Walk)
0.94 0.87
0.87 0.90
0.88 0.92
0.86 0.88
AUS-US Exchange Rate Forecast: k = 8, T = 1991Q3
RMSE Ratios (Constrained Model vs. Random Walk)
0.91 0.86
0.83 0.90
0.86 0.95
0.80 0.86
AUS-US Exchange Rate Forecast: k = 1, T = 1994Q1
RMSE Ratios (Constrained Model vs. Random Walk)
0.99 1.05
0.97 0.98
0.97 0.98
0.97 0.98
AUS-US Exchange Rate Forecast: k = 4, T = 1994Q1
RMSE Ratios (Constrained Model vs. Random Walk)
0.97 0.90
0.83 0.81
0.84 0.82
0.82 0.80
AUS-US Exchange Rate Forecast: k = 8, T = 1994Q1
RMSE Ratios (Constrained Model vs. Random Walk)
0.96 0.90
0.70 0.69
0.72 0.70
0.68 0.66
AUS-US Exchange Rate Forecast: k = 1, T = 1996Q3
RMSE Ratios (Constrained Model vs. Random Walk)
0.98 1.07
0.93 0.92
0.92 0.91
0.92 0.91
AUS-US Exchange Rate Forecast: k = 4, T = 1996Q3
RMSE Ratios (Constrained Model vs. Random Walk)
0.95 0.94
0.76 0.72
0.75 0.72
0.75 0.71
AUS-US Exchange Rate Forecast: k = 8, T = 1996Q3
RMSE Ratios (Constrained Model vs. Random Walk)
1.01 0.83
0.54 0.51
0.54 0.52
0.53 0.49
1.4
1.0
0.6
0.2
PPP
MON1
MON2
MON3


CAN-UK Exchange Rate Forecast: k = 1, T = 1991Q3
RMSE Ratios (Constrained Model vs. Random Walk)
Std Model
w/ Comm. Price
1.01 1.01
0.99 0.99
0.99 0.98
0.99 0.98
CAN-UK Exchange Rate Forecast: k = 4, T = 1991Q3
RMSE Ratios (Constrained Model vs. Random Walk)
1.11 1.09
1.00 0.98
1.00 0.98
0.98 0.97
CAN-UK Exchange Rate Forecast: k = 8, T = 1991Q3
RMSE Ratios (Constrained Model vs. Random Walk)
1.44 1.37
0.95 0.94
0.96 0.94
0.95 0.94
CAN-UK Exchange Rate Forecast: k = 1, T = 1994Q1
RMSE Ratios (Constrained Model vs. Random Walk)
1.04 1.03
1.00 0.99
1.00 0.99
0.99 0.99
CAN-UK Exchange Rate Forecast: k = 4, T = 1994Q1
RMSE Ratios (Constrained Model vs. Random Walk)
1.25 1.17
1.01 1.00
1.00 0.98
1.00 0.99
CAN-UK Exchange Rate Forecast: k = 8, T = 1994Q1
RMSE Ratios (Constrained Model vs. Random Walk)
1.49 1.40
1.02 1.02
1.01 1.01
1.02 1.02
CAN-UK Exchange Rate Forecast: k = 1, T = 1996Q3
RMSE Ratios (Constrained Model vs. Random Walk)
1.02 1.02
1.00 1.00
1.00 1.00
0.99 0.99
CAN-UK Exchange Rate Forecast: k = 4, T = 1996Q3
RMSE Ratios (Constrained Model vs. Random Walk)
1.16 1.09
1.10 1.10
1.09 1.09
1.08 1.08
CAN-UK Exchange Rate Forecast: k = 8, T = 1996Q3
RMSE Ratios (Constrained Model vs. Random Walk)
1.19 1.08
1.16 1.14
1.16 1.14
1.14 1.12
1.4
1.0
0.6
0.2
PPP
MON1
MON2
MON3



NZL-JPN Exchange Rate Forecast: k = 1, T = 1991Q3
RMSE Ratios (Constrained Model vs. Random Walk)
Std Model
w/ Comm. Price
0.98 0.92 0.99 0.96 0.98 0.95 0.99 0.96
NZL-JPN Exchange Rate Forecast: k = 4, T = 1991Q3
RMSE Ratios (Constrained Model vs. Random Walk)
Std Model
w/ Comm. Price
0.94 0.71 0.97 0.88 0.96 0.87 0.96 0.88
NZL-JPN Exchange Rate Forecast: k = 8, T = 1991Q3
RMSE Ratios (Constrained Model vs. Random Walk)
Std Model
w/ Comm. Price
0.73 0.51 0.95 0.90 0.96 0.90 0.94 0.88
NZL-JPN Exchange Rate Forecast: k = 1, T = 1994Q1
RMSE Ratios (Constrained Model vs. Random Walk)
Std Model
w/ Comm. Price
0.98 0.89 1.00 0.97 1.00 0.97 1.00 0.97
NZL-JPN Exchange Rate Forecast: k = 4, T = 1994Q1
RMSE Ratios (Constrained Model vs. Random Walk)
Std Model
w/ Comm. Price
0.93 0.45 1.00 0.67 1.00 0.66 1.00 0.65
NZL-JPN Exchange Rate Forecast: k = 8, T = 1994Q1
RMSE Ratios (Constrained Model vs. Random Walk)
Std Model
w/ Comm. Price
0.72 0.44 0.98 0.92 0.98 0.92 0.97 0.90
NZL-JPN Exchange Rate Forecast: k = 1, T = 1996Q3
RMSE Ratios (Constrained Model vs. Random Walk)
Std Model
w/ Comm. Price
0.98 0.90 1.00 0.99 1.00 0.98 1.00 0.99
NZL-JPN Exchange Rate Forecast: k = 4, T = 1996Q3
RMSE Ratios (Constrained Model vs. Random Walk)
Std Model
w/ Comm. Price
0.90 0.64 1.04 1.07 1.02 1.02 1.02 1.04
NZL-JPN Exchange Rate Forecast: k = 8, T = 1996Q3
RMSE Ratios (Constrained Model vs. Random Walk)
Std Model
w/ Comm. Price
0.69 0.46 1.00 0.99 0.97 0.96 0.97 0.95
1.4
1.0
0.6
0.2
PPP
MON1
MON2
MON3

Figure 5. Simulated Out-of-Sample Forecast Comparisons

RMSE Ratios: Models vs. Random Walk

Unconstrained Specifications with and without Commodity Prices

(See Charts on Next Three Pages)

1. The charts present ratios of root mean square forecast errors (RMSEs) produced by a model relative to those produced by the random walk. The models (PPP, Mon1, Mon2, & Mon3) are **estimated in-sample without pre-imposing any cointegrating parameter constraints**, as specified below:

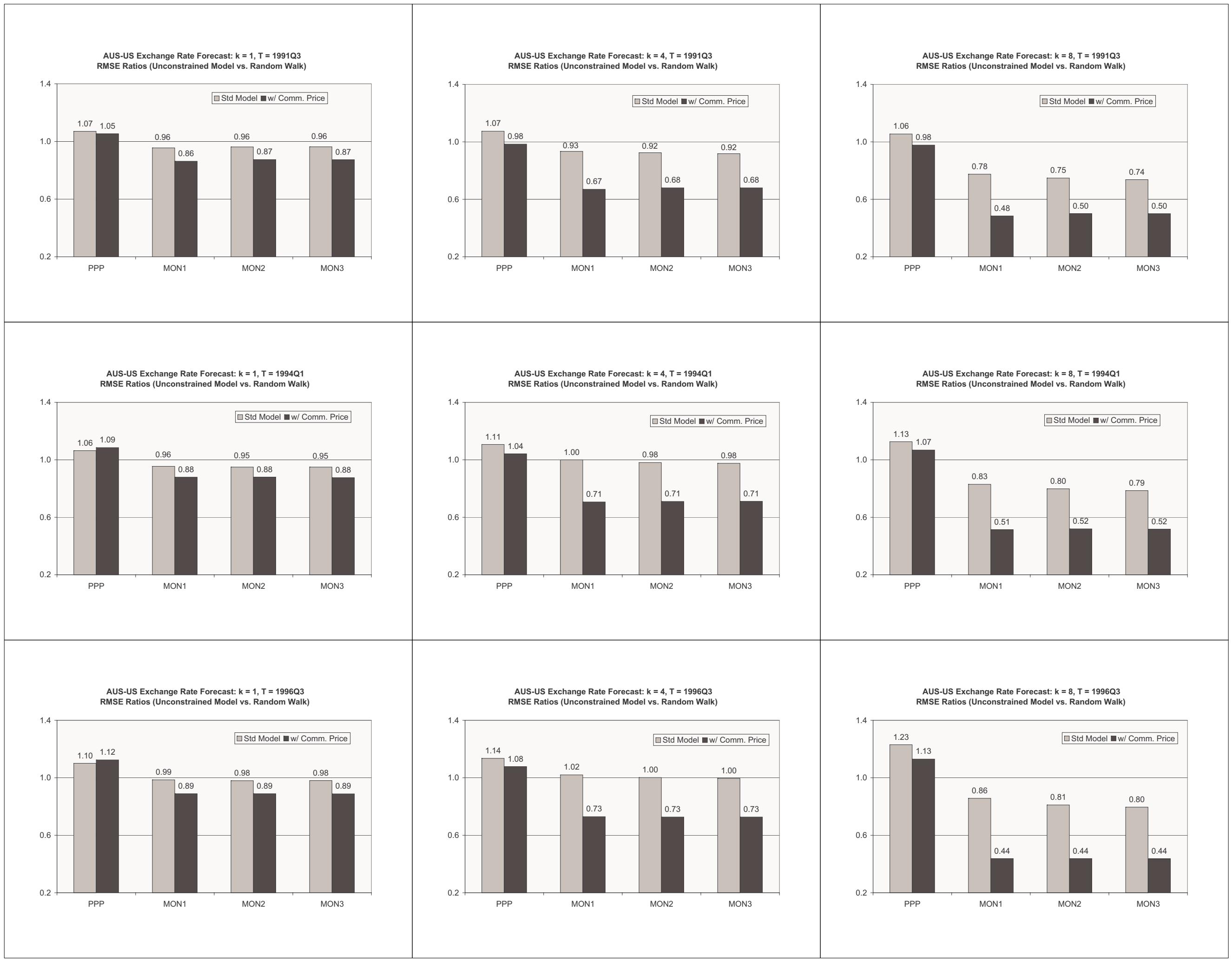
2. The forecast equations are (with and without the $\beta_{cp}p_t^{com}$ term):
   - PPP: $\Delta s_{t+k} = \alpha + \lambda_k[\beta_p (p_t^* / p_t) + \beta_{cp}p_t^{com} - s_t] + \varepsilon_{t+k}$
   - MON1: $\Delta s_{t+k} = \alpha + \lambda_k [\beta_m (m_t^* / m_t) - \beta_y (y_t^* / y_t) + \beta_{cp}p_t^{com} - s_t] + \varepsilon_{t+k}$
   - MON2: $\Delta s_{t+k} = \alpha + \lambda_k [\beta_m (m_t^* / m_t) - \beta_y (y_t^* / y_t) + \beta_i (i_t^* - i_t) + \beta_{cp}p_t^{com} - s_t] + \varepsilon_{t+k}$
   - MON3: $\Delta s_{t+k} = \alpha + \lambda_k [\beta_m (m_t^* / m_t) - \beta_y (y_t^* / y_t) - \beta_i (i_t^* - i_t) + \beta_\pi (\pi_t^* - \pi_t) + \beta_{cp}p_t^{com} - s_t] + \varepsilon_{t+k}$

   where k is the forecast horizon. All variables are in logs except for short-term interest rates and inflation rates.

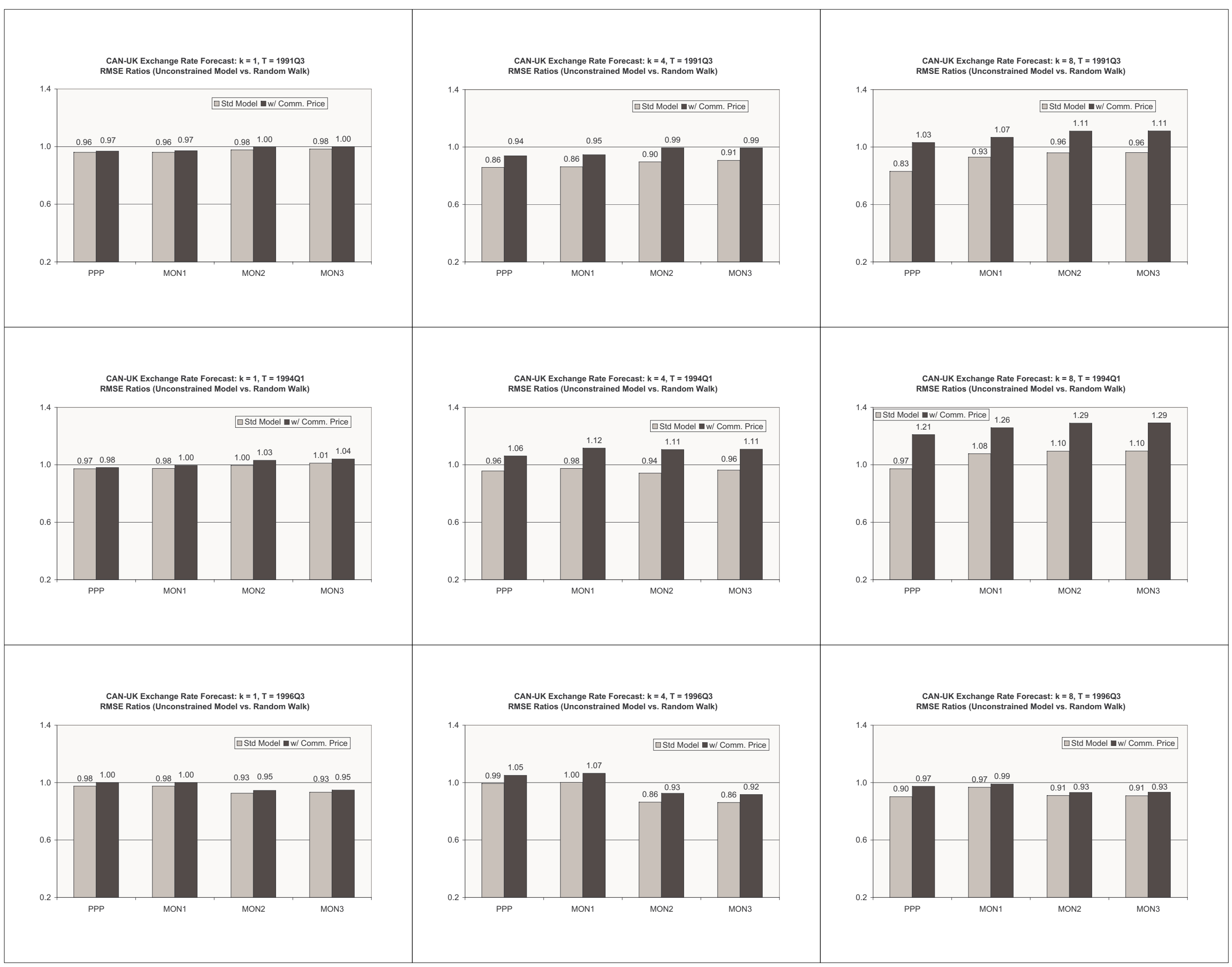
3. The lighter bars are RMSE ratios (relative to RW forecast) of forecast results from "Standard Models" without including commodity prices in the above equations; the darker bars are from models including commodity prices as specified above.

4. The nine charts on each page are organized as follows:
   a. From left to right: forecast horizon k increases from 1, 4, to 8
   b. From top to bottom: forecast period decreases from the last 40 to 30 to 20 observations. (This corresponds to forecast starting points: 1991Q3, 1994Q1, and 1996Q3.)

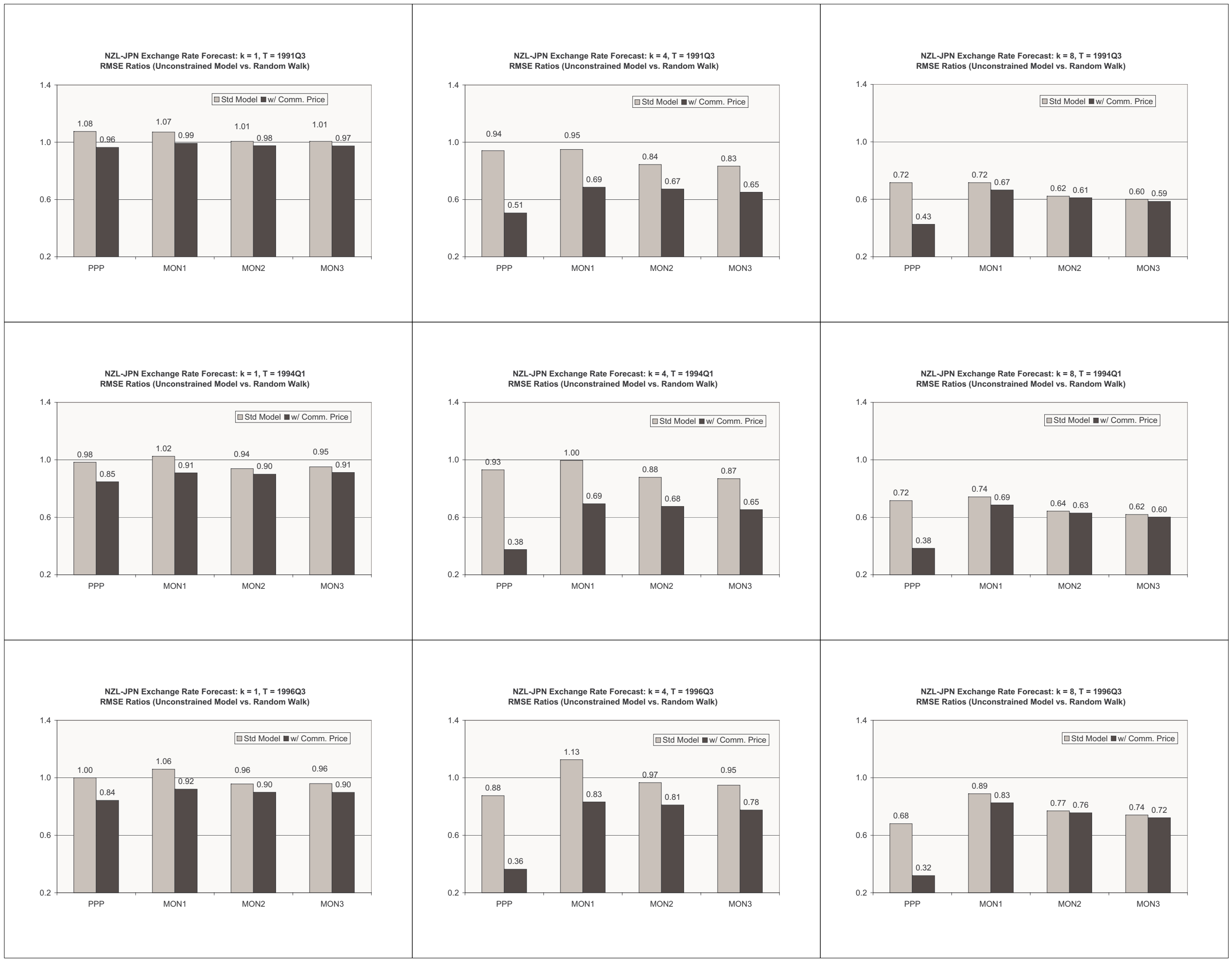
5. See texts for more discussions. Three representative currency pairs are presented here only; results for other 6 currency pairs are available upon request.



AUS-US Exchange Rate Forecast: k = 1, T = 1991Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
1.07 1.05 0.96 0.86 0.96 0.87 0.96 0.87
PPP MON1 MON2 MON3
AUS-US Exchange Rate Forecast: k = 4, T = 1991Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
1.07 0.98 0.93 0.67 0.92 0.68 0.92 0.68
PPP MON1 MON2 MON3
AUS-US Exchange Rate Forecast: k = 8, T = 1991Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
1.06 0.98 0.78 0.48 0.75 0.50 0.74 0.50
PPP MON1 MON2 MON3
AUS-US Exchange Rate Forecast: k = 1, T = 1994Q1
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
1.06 1.09 0.96 0.88 0.95 0.88 0.95 0.88
PPP MON1 MON2 MON3
AUS-US Exchange Rate Forecast: k = 4, T = 1994Q1
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
1.11 1.04 1.00 0.71 0.98 0.71 0.98 0.71
PPP MON1 MON2 MON3
AUS-US Exchange Rate Forecast: k = 8, T = 1994Q1
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
1.13 1.07 0.83 0.51 0.80 0.52 0.79 0.52
PPP MON1 MON2 MON3
AUS-US Exchange Rate Forecast: k = 1, T = 1996Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
1.10 1.12 0.99 0.89 0.98 0.89 0.98 0.89
PPP MON1 MON2 MON3
AUS-US Exchange Rate Forecast: k = 4, T = 1996Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
1.14 1.08 1.02 0.73 1.00 0.73 1.00 0.73
PPP MON1 MON2 MON3
AUS-US Exchange Rate Forecast: k = 8, T = 1996Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
1.23 1.13 0.86 0.44 0.81 0.44 0.80 0.44
PPP MON1 MON2 MON3



CAN-UK Exchange Rate Forecast: k = 1, T = 1991Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
0.96 0.97 0.96 0.97 0.98 1.00 0.98 1.00
PPP MON1 MON2 MON3
CAN-UK Exchange Rate Forecast: k = 4, T = 1991Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
0.86 0.94 0.86 0.95 0.90 0.99 0.91 0.99
PPP MON1 MON2 MON3
CAN-UK Exchange Rate Forecast: k = 8, T = 1991Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
0.83 1.03 0.93 1.07 0.96 1.11 0.96 1.11
PPP MON1 MON2 MON3
CAN-UK Exchange Rate Forecast: k = 1, T = 1994Q1
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
0.97 0.98 0.98 1.00 1.00 1.03 1.01 1.04
PPP MON1 MON2 MON3
CAN-UK Exchange Rate Forecast: k = 4, T = 1994Q1
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
0.96 1.06 0.98 1.12 0.94 1.11 0.96 1.11
PPP MON1 MON2 MON3
CAN-UK Exchange Rate Forecast: k = 8, T = 1994Q1
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
0.97 1.21 1.08 1.26 1.10 1.29 1.10 1.29
PPP MON1 MON2 MON3
CAN-UK Exchange Rate Forecast: k = 1, T = 1996Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
0.98 1.00 0.98 1.00 0.93 0.95 0.93 0.95
PPP MON1 MON2 MON3
CAN-UK Exchange Rate Forecast: k = 4, T = 1996Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
0.99 1.05 1.00 1.07 0.86 0.93 0.86 0.92
PPP MON1 MON2 MON3
CAN-UK Exchange Rate Forecast: k = 8, T = 1996Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model w/ Comm. Price
0.90 0.97 0.97 0.99 0.91 0.93 0.91 0.93
PPP MON1 MON2 MON3


NZL-JPN Exchange Rate Forecast: k = 1, T = 1991Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
Std Model
w/ Comm. Price
1.4
1.0
0.6
0.2
PPP
MON1
MON2
MON3
1.08 0.96
1.07 0.99
1.01 0.98
1.01 0.97
NZL-JPN Exchange Rate Forecast: k = 4, T = 1991Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
0.94 0.51
0.95 0.69
0.84 0.67
0.83 0.65
NZL-JPN Exchange Rate Forecast: k = 8, T = 1991Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
0.72 0.43
0.72 0.67
0.62 0.61
0.60 0.59
NZL-JPN Exchange Rate Forecast: k = 1, T = 1994Q1
RMSE Ratios (Unconstrained Model vs. Random Walk)
0.98 0.85
1.02 0.91
0.94 0.90
0.95 0.91
NZL-JPN Exchange Rate Forecast: k = 4, T = 1994Q1
RMSE Ratios (Unconstrained Model vs. Random Walk)
0.93 0.38
1.00 0.69
0.88 0.68
0.87 0.65
NZL-JPN Exchange Rate Forecast: k = 8, T = 1994Q1
RMSE Ratios (Unconstrained Model vs. Random Walk)
0.72 0.38
0.74 0.69
0.64 0.63
0.62 0.60
NZL-JPN Exchange Rate Forecast: k = 1, T = 1996Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
1.00 0.84
1.06 0.92
0.96 0.90
0.96 0.90
NZL-JPN Exchange Rate Forecast: k = 4, T = 1996Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
0.88 0.36
1.13 0.83
0.97 0.81
0.95 0.78
NZL-JPN Exchange Rate Forecast: k = 8, T = 1996Q3
RMSE Ratios (Unconstrained Model vs. Random Walk)
0.68 0.32
0.89 0.83
0.77 0.76
0.74 0.72