Commodity Currencies and Empirical Exchange Rate Puzzles

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Abstract

This paper re-examines empirical exchange rate puzzles by focusing on three OECD economies (Australia, Canada, and New Zealand) where primary commodities constitute a significant share of their exports. For Australia and New Zealand especially, we find that the U.S. dollar price of their commodity exports (generally exogenous to these small economies) has a strong and stable influence on their floating real rates, with the quantitative magnitude of the effects consistent with predictions of standard theoretical models. However, after controlling for commodity price shocks, there is still a PPP puzzle in the residual. Nevertheless, the results here are relevant to many developing country commodity exporters, as they liberalize their capital markets and move towards floating exchange rates.

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1. Introduction

The connection between economic fundamentals and exchange rate behavior has been one of the most controversial issues in international finance, manifesting itself in various major empirical puzzles such as the Meese-Rogoff (1983) puzzle and the purchasing power parity (PPP) puzzle (Rogoff 1996).[1] Recent research efforts to confront these challenges have explored new approaches in both theoretical and empirical fronts, including incorporating non-linearity in modeling exchange rate dynamics.[2] Alternatively, it has also been recognized that if one could find a real shock that were sufficiently volatile, one could potentially go a long ways towards resolving these major empirical exchange rate puzzles. For most OECD economies, however, it is hard to know what that shock might be, much less measure it.[3] In this paper, we focus on three OECD economies where a potential dominant real shock may be identified, and explore how controlling for this real shock may help shed light on empirical exchange rate puzzles.

In Canada, Australia and New Zealand, because primary commodities constitute a significant component of their exports, world commodity price fluctuations – generally exogenous to these small countries for all but a few goods – potentially explain a major component of their terms of trade fluctuations.[4] In fact, researchers at the Bank of Canada have claimed for many years that not only do their empirical exchange rate equations fit out-of-sample, one can even use variants to successfully predict the exchange rate, both unconditionally and in response to policy alternatives.[5] A key element of the Canadian equation involves augmenting the standard model by a terms of trade variable reflecting the volatile movements in world prices of Canadian commodity exports, particularly non-energy commodities. Researchers at the Reserve Bank of Australia have at times been even more ebullient, finding that over the

[1] See Frankel and Rose (1995) and Froot and Rogoff (1995) for a comprehensive survey of the empirical research on exchange rates.
[2] Examples of recent papers that explore non-linear exchange rate responses to deviations from economic fundamentals include Taylor and Peel (2000) and Taylor (2001).
[3] Oil prices certainly have the volatility and there is some evidence that they impact the terms of trade (Backus and Crucini, 2000), but adding these variables to standard monetary equations does not seem to do the trick.
[4] Simply incorporating standard measures of terms of trade as an explanatory variable would not be meaningful for most OECD countries, since the terms of trade contain a large component that moves mechanically with the exchange rate (see Obstfeld and Rogoff, 2000). This is most likely due to a combination of wage and price rigidities interacting with partial pass-through. Our main explanatory variable here is not the terms of trade but indices of world commodity prices, which presumably do not move automatically with these small countries exchange rates. See section 4 for further discussion.

1990's, one could have earned a substantial excess profit in trading on the Australian dollar by properly incorporating forecastable terms of trade movements into exchange rate forecasts.[6] Finally, a similar framework has also been extended to the New Zealand dollar.[7]

Our paper aims to address the following two questions. Is it true that commodity price shocks explain a significant share of exchange rate movements for these currencies? And if so, does the introduction of commodity prices more broadly solve the PPP puzzle, opening the door to salvation of standard monetary exchange rate models for these currencies? Affirmative answers here might encourage researchers to try harder to search for corresponding real shocks to the major currencies. More broadly, from a policy stand-point, understanding the effects of commodity price shocks on exchange rates is of considerable interest to developing countries, particularly as they liberalize capital market controls and adopt more flexible exchange rates. If commodity prices can indeed be shown to be a consistent and empirically reliable factor in empirical exchange rate equations, the finding would have important implications across a variety of policy issues, not least concerning questions such as how to implement inflation-targeting in developing countries.[8]

The outline of this paper is as follows. In section 2 of the paper, we give a brief overview of the economic environment in the three countries, and provide some simple, but striking, evidence on just how closely movements in these currencies seem to track the corresponding world price of their commodity exports. The strong correlations come through not only for cross rates against the U.S. dollar, but also when the British pound and a broad index of non-US-dollar currencies are used as the numeraire. In section 3, we go on to see whether the visual evidence stands up to closer scrutiny, focusing on simple empirical models, not least because the data sample is limited and richer dynamic models would lack credibility. We find that for New Zealand and Australia, the connection between commodity prices and exchange rates holds up remarkably well, and appears quite robust to alternative assumptions on the underlying time

[5] See, for example, Amano and van Norden (1993) and Djoudad, Murray, Chan and Dow (2001).
[6] See, for example, Gruen and Kortian (1996).
[7] See again Djoudad, Murray, Chan and Dow (2001).
[8] There has been related work for developing countries that looks at cross-country panel data. For example, Mendoza (1995) finds that terms of trade shocks account for a significant portion of variation in output in developing countries, whereas Bidarkota and Crucini (2000) find a strong connection between shocks to commodity prices and the terms of trade.

series properties. Though there is some evidence of structural breaks, especially at the time these countries switched to formal inflation targeting, the general size of the contemporaneous correlation between commodity prices and exchange rates nevertheless seems relatively consistent, both across time and across countries. The commodity price elasticity estimate is typically in the neighborhood of 0.75 for both Australia and New Zealand. For Canada, the evidence is more mixed, with the correlation between exchange rates and commodity prices much more sensitive to detrending. We go on in section 4 to consider two potential forms of misspecifications. We first ask whether the relationship might result from the countries having market power in their commodity exports, and find that this is not the case. In addition, we use commodity price indices as instruments for standard measures of terms of trade, and conclude that world commodity prices in fact better capture exogenous shocks to these countries' terms of trade than standard measures do. Section 5 then offers a structural interpretation of the estimates in light of standard exchange rate models. We argue that the quantitative size of our estimates is quite plausible.

Having found a robust connection between commodity export prices and exchange rates, we extend the analysis in section 6 to ask whether the inclusion of commodity prices can help provide stronger empirical support for canonical exchange rate models. By controlling for this major source of real shocks, one might hope the standard exchange rate equations - adjusted for commodity price shocks - might perform better for "the commodity currencies" than they have been found to perform for the major currencies. However, our results do not offer very strong encouragement for this point of view. In fact, we show that standard monetary variables are unlikely to work well in explaining exchange rate behavior, at least in linear models, because real exchange rates remain extremely persistent. The final section concludes.

2. Background and Graphical Evidence

2.1. Background

To effectively explore the temporal relationship between exchange rate behavior and commodity price shocks, we focus on developed economies where internal and external markets operate with little intervention, and where floating exchange rate regimes have been implemented for a sufficiently long period of time.[9] From a macroeconomic perspective, Australia, Canada, and New Zealand are near perfect examples of such well-developed, small open economies. All three are highly integrated into global capital markets and are active participants in international trade. And in terms of monetary and exchange rate policies, they have all been operating under a flexible exchange rate regime for well over a decade. Canada began floating its currency before the collapse of Bretton Woods, in 1970. Australia and New Zealand abandoned their exchange rate pegs in 1983 and 1985 respectively, as part of the economic reform efforts to revitalize their domestic economies. Moreover, around 1990, all three adopted some variant of inflation targeting monetary policy.[10]

To varying degrees, all three countries can plausibly be described as "commodity economies", because of the large share primary commodities occupy in their production and exports. For at least the past decade, commodities have maintained a 60% share of Australia's total exports, with wool, wheat, and various metals examples of its leading exports. In New Zealand, while the share has declined from a hefty two-thirds in the late 1980s, commodities continue to account for more than half of its total exports in recent years. By comparison, Canada has a larger and more developed industrial base, but still, it continues to rely more than a quarter of its exports on commodities such as base metals, forestry products, and crude oil. Despite the relatively small size of their overall economies, these countries retain a significant share of the global market for a few of their export products. In New Zealand, for instance, 46 million sheep cohabit with 3.8 million people. Not surprisingly, only 20 percent of its meat production is consumed domestically, and New Zealand supplies close to half of the total world exports of lamb and mutton. Canada similarly dominates the world market in forestry products, and Australia holds significant shares of the global exports in wool and iron ore. However, while each country may have some market power for a few key goods, they are, on the whole, price takers in world markets for the vast majority of their commodity exports.

[9] Among other OECD countries, Finland, Norway, and the United Kingdom also export significant amount of primary commodities (e.g. forestry products for Finland and North Sea Oil for the latter two countries). Finland and Norway are excluded in our study as they operated under regulated exchange regimes for much of the past two decades. The United Kingdom has about a fifth of its exports in crude oil, but we find little connection between its exchange rate behavior and the world price of oil.

[10] We refer the reader to Zettelmeyer (2000) for a more thorough discussion of the conduct of monetary policy in these three countries in the 1990s.

2.2. Graphical Evidence and Data Description

Figures 1b through d show the value of Australian dollar relative to three reference currencies – the U.S. dollar, the British Pound, and a non-US-dollar currency basket – plotted alongside the world price of Australia's major non-energy commodities. As a means of comparison, Figure 1a plots the Australian-US real exchange rate with a couple standard macroeconomic variables: the real income differentials and the real interest rate differentials between the two countries. The corresponding graphs for Canada and New Zealand are shown in Figures 2 and 3.

For all three countries, the sample period starts shortly after the float of the particular home currency.[11] The real exchange rates in these graphs (and in all the subsequent analyses) are end-of-quarter nominal rates, expressed as the foreign exchange values of the domestic currency, adjusted by the relative CPIs. An increase in the real rates thus represents a rise in the relative price of home goods or, a real appreciation for the home country. The non-dollar basket is adopted from the Broad Index of the Federal Reserve. It is a composite of over 30 non-US-dollar currencies, covering all major trading partners of the United States each weighted by their respective trade shares.[12] By measuring the relevant home currencies against different anchors, especially the Broad Index covering many developing countries, we hope to insulate our analysis from being driven by shocks to the U.S. economy and the movements in the U.S. dollar.

The country-specific commodity price indices cover non-energy commodities only, and are geometric averages of the world market prices of the major commodities produced in each country, weighted by their corresponding domestic production share.[13] Individual real commodity prices are quarterly averages of the world market transaction prices in U.S. dollars,

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[11] To abstract away from the issue of non-stationarity in nominal prices, we focus on real exchange rate behavior in this paper. See Section 3 for a more detailed discussion.

[12] It does not matter that the home-country currency appears in our non-dollar index (the New Zealand and Australia weights are zero/very small), since it essentially factors out when we construct the exchange rate of the non-dollar index against the home currency. There is no particular significance to using U.S. trade weights in our analysis; we adopt the broad index of the Federal Reserve as a convenient check for the robustness of our results.

[13] We separate out energy and non-energy commodities because none of these countries is a clear large net exporter of energy commodities, in contrast with non-energy commodities. Including an energy commodity price index does not change any of the results for Australia and New Zealand, and for Canada, while it has some explanatory power for movements in the US-Canada rate, the results are not robust to different anchor currencies.

deflated by the U.S. CPI. The commodities included in each index and their corresponding weights are listed in the data appendix.

Looking at these sets of graphs, three features especially stand out. First, whereas the two standard monetary exchange rate variables – interest rate and output differentials – exhibit no obvious visual correlation with the exchange rates, the correlations between commodity prices and various exchange rates are strikingly apparent. The commodity price and exchange rate series not only appear to mirror each other in movement, the magnitude of their swings are also similar. (This observation is confirmed by the contemporaneous regression results presented in Table A.1 of the Appendix. Comparing across various currency pairings, we note the remarkable similarity in the coefficient estimates.) Secondly, these real exchange rates appear highly persistent and possibly non-stationary, a point we will address in more details in Section 3. Lastly, the well-documented long-term decline of global commodity prices seems clearly reflected in these country-specific series as well. In the following section, we explore further just how strong and robust the apparent correlations are, and the role common trends (stochastic or not) may play in explaining the co-movements of real exchange rates and commodity prices.

3. Empirical Analysis

While establishing simple correlations seems an appropriate starting point in light of earlier empirical failures, formal empirical analysis cannot avoid addressing the issue of how best to model a small sample of data with near unit root behavior. Our short samples of fewer than 100 quarterly observations simply preclude any meaningful test of stationarity, a well-documented problem that has stimulated numerous innovative studies using long-horizon time series or panel data, coupled with various econometric techniques. In this paper, we rely on the considerable empirical evidence suggesting that real exchange rates are stationary, possibly with a trend.[14] For example, Froot and Rogoff (1995) present evidence that the half-life of real exchange rate shocks in linear models is roughly 3-4 years across a wide variety of historical data. Culver and Papell (1999) and Wu (1996), among others, show mean-reversion in the post-

[14] The same is arguably true for commodity prices. See Borensztein and Reinhart (1994); Bleaney (1996); and Cashin, Liang and McDermott (2000).

Bretton Woods real exchange rates of most industrialized countries.[15] In addition, using a century of annual data, Bleaney (1996) demonstrated that the trade-weighted Australian real exchange rate, along with the world price of primary commodities (relative to that of manufacturer), are both trend-stationary.

Ruling out non-stationarity/stochastic trends a priori, we mainly focus on the case where real exchange rates and real commodity prices are treated as stationary, possibly with trends.[16] However, in the first subsection below, we consider several alternative underlying data-generating processes, including I(1) processes, as robustness checks for our results. We find that for Australia and New Zealand, the connection between real exchange rates and the world price of their commodity exports is quite strong and stable (whether or not we exclude unit roots), while that for the Canadian dollar seems less robust, especially to detrending. We then examine the stability of these parameter estimates in Section 3.2.

3.1. Trends, Serial Correlations, and Non-Stationarity

We first present estimates for the commodity price elasticity of real exchange rate for the three countries, treating both series as stationary with a linear trend (see figures 4-6 for the linearly detrended series). The OLS coefficient estimates are reported in the first column of Tables 1a-c below. Since results based on different anchor currencies are similar, to conserve space, only results for the U.S. dollar rates are reported here. For Australia and New Zealand, we note that the elasticity estimates show up slightly higher but in general consistent with those obtained without the time trend (see Table A.1 in the appendix). From the second and the last columns, we see that these estimates also appear robust to alternative detrending methods: Hodrick-Prescott filtering and first differencing.[17] For Canada, the positive correlation between

[15] One might argue that even though most real exchange rates appear to be stationary, the commodity currencies might be an exception if commodity prices themselves have a unit root, which as noted in the previous footnote, does not appear to be the case. Even if commodity prices do have a unit root, it will not necessarily be the case that real exchange rates do too. Over the very long run, countries can substitute out of commodity production into manufactures if the relative price of commodities drifts too low. Korea today exports primarily manufactured goods, but in 1960 almost 90% of exports were commodities.

[16] The time trend in the real exchange rate could reflect, for example, deterministic evolutions in the sectoral productivity differences of the Balassa-Samuelson model (see Sections 5 and 6).

[17] We recognize the limitation and potential problems associated with each of these filters. That is, both linear and HP filters will result in spurious cycles should the underlying series be difference stationary. While the HP filter does not assume a stable trend process over time, it suffers from the well-documented end-point problem. The first

commodity price and exchange rate does not seem to survive detrending, an issue we will discuss further.

Table 1: Real Exchange Rates and Commodity Prices:

Different Assumptions on the Data Generating Processes

1a: Australia

Dependent Variable: Log of Real Exchange Rate, vs. U.S. Dollar

Sample Period: 1984Q1 – 2001Q2

	I(0)/Deterministic Trends			I(1)/Stochastic Trends	
	Linear Trend + Newey-West S.E. [1]	HP Filter + Newey-West S.E. [2]	Linear Trend + AR(1) Residuals [3]	Cointegration: Dynamic OLS [4]	Non-Cointegration: 1st Differencing [5]
Ln(Real Commodity Prices)	0.81* (0.12)	0.58 * (0.12)	0.54 * (0.14)	0.39 * (t = 6.19)	0.47 * (0.14)
AR(1) root			0.88 * (0.05)		
Durbin-Watson stat	0.36		2.15		
Adj. R^2	0.57	0.37	0.86	0.36	0.07
N Obs.	70				

Note: * indicates significance at the 5% level. Newey-West heteroskedasticity and autocorrelation consistent (HAC) standard errors in parentheses (except for the AR(1) specification).

1.) $\ln(\text{Real Exchange Rate})_t = \alpha + \beta^* t + \gamma^* \ln(\text{Real Commodity Price})_t + \varepsilon_t$

using non-parametric GMM Newey-West approach to correct for the biased standard errors estimate.

2.) Hodrick-Prescott Detrended $\ln(\text{Real Exchange Rate})_t = \alpha + \beta^*$ HP Detrended $\ln(\text{Real Commodity Price})_t + \varepsilon_t$, using non-parametric GMM Newey-West approach to correct for the biased standard errors estimate.

3.) $\ln(\text{Real Exchange Rate})_t = \alpha + \beta^* t + \gamma^* \ln(\text{Real Commodity Price})_t + \varepsilon_t$, where ε_t follows an AR(1)

4.) $\ln(\text{Real Exchange Rate})_t = \alpha + \gamma^* \ln(\text{Real Commodity Price})_t + \gamma_1^* \Delta\ln(\text{Real Commodity Price})_{t+1} + \gamma_0^* \Delta\ln(\text{Real Commodity Price})_t + \gamma_{-1}^* \Delta\ln(\text{Real Commodity Price})_{t-1} + \varepsilon_t$, where Δ is the first difference operator. Here real exchange rates and real commodity prices are assumed to be non-stationary, so DOLS procedure is used to obtain super-consistent estimators for the cointegrating vector (Stock and Watson 1993). T-ratios, which are asymptotically standard normal, are reported.

5.) $\Delta\ln(\text{Real Exchange Rate})_t = \alpha + \beta^* \Delta\ln(\text{Real Commodity Price})_t + \varepsilon_t$, using non-parametric GMM Newey-West approach to correct for potential serial correlation in the standard errors estimate

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difference filter removes any long-term trend but also potentially important information contained in the levels of the series; it is mainly relevant when commodity prices and real exchange rates are non-cointegrated I(1) series.

Table 1 (Continued)

1b: Canada

Dependent Variable: Log of Real Exchange Rate, vs. U.S. Dollar

Sample Period: 1973Q1 - 2001Q2

	I(0)/Deterministic Trends			I(1)/Stochastic Trends	
	Linear Trend + Newey-West S.E. [1]	HP Filter + Newey-West S.E. [2]	Linear Trend + AR(1) Residuals [3]	Cointegration: Dynamic OLS [4]	Non-Cointegration: 1st Differencing [5]
Ln(Real Commodity Prices)	0.21 (0.15)	0.09 (0.07)	0.04 (0.07)	0.40 * (t = 11.94)	0.05 (0.06)
AR(1) root			0.96 * (0.03)		
Durbin-Watson stat	0.10		1.92		
Adj. R^2	0.06	0.03	0.96	0.56	-0.00
N Obs.	114				

1c: New Zealand

Dependent Variable: Log of Real Exchange Rate, vs. U.S. Dollar

Sample Period: 1986Q1 - 2001Q2

	I(0)/Deterministic Trends			I(1)/Stochastic Trends	
	Linear Trend + Newey-West S.E. [1]	HP Filter + Newey-West S.E. [2]	Linear Trend + AR(1) Residuals [3]	Cointegration: Dynamic OLS [4]	Non-Cointegration: 1st Differencing [5]
Ln(Real Commodity Prices)	1.10 * (0.32)	0.73 * (0.20)	0.51 * (0.21)	0.58 * (t = 6.17)	0.59 * (0.26)
AR(1) root			0.95 * (0.05)		
Durbin-Watson stat	0.19		1.99		
Adj. R^2	0.37	0.25	0.90	0.40	0.10
N Obs.	62				

The Durbin-Watson statistics in these regressions indicate that significant positive serial correlations remain in the residuals, even after detrending. Leaving its economic implication to Section 6, here we address alternative methods for correcting the biased standard errors estimates. For the majority of the analysis in this paper, the kernel-based nonparametric GMM estimator of Newey-West (1987) is used to account for the serial correlation. However, because such non-parametric estimators have poor small sample properties, the third columns in Tables 1a-c present estimation results from an alternative parametric specification, where the error terms are assumed to follow a first order autoregressive process. The AR(1) specifications effectively bring the Durbin-Watson statistics back towards 2, and while lowering the coefficients slightly, still give estimates consistent with earlier findings.[18]

As already noted, tests of unit roots or cointegration have little statistical power in short time series. In fact, Blough (1992), Cochrane (1991), and Faust (1996) contend that in finite samples, a stationary process can always be arbitrarily well approximated by a non-stationary process (and vice versa).[19] Following up on this observational equivalence idea, Engel (2000) further argues that the rejections of unit-root null in long horizon real exchange rate data may be the result of size distortions. If our real exchange rate and commodity price series are indeed non-stationary, the estimates and significance tests performed so far, based on classical statistical methods, would be invalid. Therefore, a robustness check we consider next takes up the alternative assumption that real exchange rates and real commodity prices follow unit root processes.

If the real exchange rate and commodity price series are non-stationary but *not* cointegrated, the estimation needs to be done in first-differences to avoid spurious regression. We have already seen that the first differenced series produce similar estimates as under the linear-trend specification (last columns of Table 1a-c). However, if we instead assume that the two series *are* cointegrated, the first-differencing approach would no longer be appropriate. The 4th columns in Tables 1a-c present results from dynamic OLS (DOLS) of Stock and Watson

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[18] In addition, the autocorrelation coefficient estimates, in the range of 0.88 to 0.96, are broadly similar to what we see in PPP regressions (See Froot and Rogoff 1995). We will come back to these estimates in Section 6.
[19] See Engel (2000) and Mark (2001), for example, for more discussion on this observational equivalence issue.

(1993), a specification designed to estimate cointegrating relations.[20] The DOLS approach produces the correct standard errors for the "superconsistent" point estimates of the cointegrating vectors. Importantly, these estimates are robust to the potential endogeneity of commodity prices, an issue we discuss in Section 4.[21] Our small sample sizes notwithstanding, the elasticity estimates for Australia and New Zealand from DOLS are not far off from those obtained under the assumption of stationarity, and still produce estimates significantly different from zero.[22] As the coefficient estimates for Australia and New Zealand appear robust to various assumptions on the underlying data generating processes, we will proceed with the linear trend model, in accordance with our view that these series are trend-stationary.

For Canada, however, unlike in the trend-stationary models, commodity price shows up as significant under DOLS, likely reflecting common trends. We note that one cannot entirely dismiss the significance of the common long-term trend between Canadian dollar and its commodity export prices. Indeed, the downward drift in both series may be intimately connected; we simply cannot statistically demonstrate any such connection here. On the other hand, this elusive correlation may reflect Canada's ambiguous status as a "true commodity economy". After all, commodities are the minority in its export base, especially compared to the case of New Zealand and Australia.[23] Structural breaks occurring somewhere over the thirty-year period is certainly another possibility. Indeed, looking at the Canadian-U.S. rate post-1985 only (a sample period comparable to those used for Australia and New Zealand), we obtain significant positive coefficient estimates of around 0.3 under both the linear and the hp filters. However, unlike the robustness we observed in the Australian and New Zealand estimates, the significant correlation disappears when the Canadian rate is measured relative to other anchor currencies.

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[20] We are aware of the inference problem put forth by Elliott (1998) that applying cointegration methods on local-to-unit root processes may introduce biases and inefficiencies. Here, we are simply using it as a robustness check.
[21] See Hamilton (1994), Ch. 19.
[22] While asymptotically, the Stock-Watson dynamic OLS procedure yields the same parameter estimates as conventional OLS, this may not be the case in finite samples. We experimented with longer lead and lag lengths, and the estimation results are similar.
[23] It is important to emphasize that we focus here on non-energy commodities only. In gross terms at least, Canada is a significant exporter of coal and natural gas. As mentioned in footnote 13, our explorations with the inclusion of energy prices separately did not significantly change our results for the Canadian dollar, but the issue needs to be investigated more thoroughly than we take up here.

3.2. Parameter Stability

As mentioned earlier, all three countries adopted inflation targeting policy in the early 1990s. Together with our findings in Section 3.1 that the estimates for the Canadian dollar appear qualitatively different when we look at a shorter sample, testing for possible structural breaks seems warranted. Table II below present results from the classic Chow test on pre-selected potential breakpoints and the Hansen (1992) test for structural break of unknown timing. As we are interested in possible instability in the commodity price elasticities, but not so much in shifts in underlying time trends, we use HP-filtered variables for this analysis.

Table II: Representative Parameter Stability Tests under HP Filter

Dependent Variable: Log of Real Exchange Rate, vs. U.S. Dollar

Detrended ln(Real Exchange Rate)$_t$ = $\alpha + d_t + (\beta + \gamma{*}d_t)$* Detrended ln(Real Commodity Price)$_t$ + ε_t,

where $d_t = 1$ if $t \geq$ Breakpoint; $d_t = 0$ otherwise

	Australia		Canada		New Zealand	
	OLS + Hansen Test	OLS + Break Dummy	OLS + Hansen Test	OLS + Break Dummy	OLS + Hansen Test	OLS + Break Dummy
Ln(Real Commodity Prices): β	0.58 * (0.12)	0.60 * (0.15)	0.09 (0.07)	0.13 (0.09)	0.72 * (0.20)	1.23 * (0.45)
Dummy* Ln(Real Commodity Prices): γ		-0.05 (0.26)		-0.09 (0.13)		-0.66 (0.50)
Breakpoint tested [1]	NA	1993Q1	NA	1991Q1	NA	1990Q1
Hansen test [2]	0.49 *		0.38		0.47*	
Adj. R^2	0.37	0.36	0.03	0.02	0.25	0.26
N Obs.	70		114		62	
Sample Period	1984Q1 - 2001Q2		1973Q1 - 2001Q2		1986Q1 - 2001Q2	

Note: * indicates significance at the 5% level. Newey-West heteroskedasticity and autocorrelation consistent (HAC) standard errors in parentheses.
1. The breakpoint is the starting year for the use of formal inflation targets in the country.
2. The 5% asymptotic critical value for the Hansen individual parameter test is 0.47 (see Hansen 1992, Table 1).

As discussed in Hansen (1992), because pre-selected candidate breakpoints are often endogenous, the Chow test is likely to falsely indicate a break when none in fact exists. In our analysis, for example, the candidate break-dates are chosen to be the year each of these countries adopted formal inflation targets (1990 for New Zealand, 1991 for Canada, and 1993 for Australia). It is easy to make a case that these regime shifts were endogenous. Nevertheless, the coefficients on the time dummies in the Chow test provide little indication of parameter shifts pre- and post-inflation targeting, despite a likely bias towards doing so. The Hansen (1992) procedure is approximately the Lagrange multiplier test for the null of constant parameters, against the alternative of structural breaks of unknown timing and/or random walk parameters.[24] It similarly does not provide strong indication of parameter instability over the full sample periods.[25] We note that the Hansen test relies on asymptotic properties our small sample size may not adequately satisfy. Nevertheless, the conclusion we draw from these tests is that while there may have been some parameter shifts over time, the basic sign and magnitude of the coefficients are notably stable for this kind of data.

Given the stability of the elasticity estimates, it is natural to explore the out-of-sample forecast performance of these contemporaneous correlations, especially for Australia and New Zealand.[26] Simple rolling forecast regressions in the spirit of the original Meese-Rogoff (1983) are conducted across different forecast horizons and base currencies to predict the levels of exchange rate. In Table A.3 of the Appendix, we present ratios of the root mean squared forecast errors between a commodity price-augmented exchange rate equation and the Random Walk. We do not observe significant out-of-sample gains over the Random Walk specification, especially across benchmark currencies. However, we note that the usefulness of commodity prices in exchange rate forecasts, even for these commodity currencies, requires much more in-

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[24] This particular version of the Hansen test does require stationary regressors, or else a different distributional theory applies. So these results are valid only under the assumption that our series are trend-stationary.

[25] We also performed the same tests using linearly detrended data (see Table A.2. in the Appendix) and against different base currencies. For Australian dollar against the non-dollar basket, neither the Chow break point test nor the Hansen test shows parameter shifts. However, both tests reject parameter stability for the Australian-UK exchange rate. Estimates for Canada remain negative and insignificant for the other two anchor currencies. For the New Zealand dollar against the non-dollar basket or the British pound, neither the Chow nor the Hansen test indicates parameter instability. However, we note that the Chow test and the Hansen test do not always produce consistent conclusions. As shown in the last two columns of Table A.2, the Chow test indicates a break in 1990Q1 for linearly-detrended New Zealand data, but the Hansen test fails to reject parameter shift over the whole sample range.

[26] It is evident from the figures 1-6 that although commodity prices and exchange rates have been remarkably correlated over the sample, the relationship notably breaks down over the past two years, especially for the U.S. dollar cross rates. Thus, any out-of-sample test, particularly one that focuses on the last few years, is likely to lead to fairly negative results. We acknowledge this, but given the striking correlation over the longer sample – far more striking than for typical empirical exchange rate equations – we will focus here on a simple forecast specification. The reader should interpret our results accordingly.

depth analysis than our simple specification here offers. In particular, as pointed out by Diebold and Killian (2000), unit-root formulations, such as the vector error correction framework, may be a more appropriate specification for forecasting regardless of the true nature of the data generating processes.[27] In addition, the statistical significance of any apparent forecast improvements would need to be evaluated.[28]

4. Some Possible Misspecifications

4.1. Endogeneity of Commodity Prices

We have thus far treated commodity prices as exogenous in the exchange rate equation. In this section, we consider two possible channels of endogeneity that could potentially bias our estimates, and show that neither is likely to be dominating our results. First, as alluded to earlier, omitted variables related to cycles and shocks in the United States, or even the global economy, might affect both the exchange rate and the commodity markets independently. For example, a boom in the U.S. economy is likely to affect *all* dollar cross rates as well as the world price of the commodities Australia exports.[29] However, a boom in the U.S. would seem unlikely to have a first order impact on the Australian cross rates against the British pound or a broad set of non-U.S. dollar currencies. Yet in these non-dollar regressions, we observe similar coefficients for the commodity price variable (see, for example, Table A.1 in the Appendix). One also has to allow for the possibility of a broad boom that affects all industrial countries (except Australia), drives up commodity prices, and simultaneously exerts an independent effect on the Australian exchange rate. We note, however, that most models would predict that this independent effect (from high world growth relative to Australian growth) should tend to *depreciate* rather than appreciate the Australian currency. So, the fact that our coefficient estimates are consistently positive and of similar magnitudes across currency pairings tends to mitigate against this source of bias.[30]

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[27] The issue of forecast performance is investigated more fully in Chen (2002), though the paper focuses on nominal exchange rates.
[28] See, for example, Kilian (1997) and Clark and McCracken (2001).
[29] See Borensztein and Reinhart (1994)
[30] Canada is the exception in our analysis. Although the Canadian dollar regressions occasionally give significant negative estimates, mostly the coefficients are insignificantly different from zero. Overall, we don't observe any consistent pattern in our analysis of Canadian dollar.

A second source of endogeneity can operate through any market power these countries may hold in commodity markets. For instance, since New Zealand controls a near majority of the global sheep market, the world price of sheep may be significantly influenced by the value of New Zealand dollar. To address this potential form of endogeneity, we use a price index that incorporates *all* non-energy commodities, each weighted by their global export earning shares, as an instrument for the country-specific production-weighted price index that we have been using.[31]

Table III reports three representative results comparing GMM-IV regressions, using world price of all commodities as instruments, with their uninstrumented OLS counterparts. We employ a GMM procedure to optimally weigh the orthogonality conditions and automatically correct the standard errors for serial correlation. As evident from the table, the world commodity price series works well as an instrument for the country specific commodity prices, and the IV estimations corroborate the least-squares findings. Namely, for Australia and New Zealand, world commodity price movements are associated with large and significant real exchange rate responses, while the effects are much smaller and mostly insignificant for Canada.[32]

[31] We want to reiterate the point that despite having significant market power in a few commodities, these three countries are relatively small in the *overall* global commodity market. In 1999, for example, Australia represents less than 5 percent of the total world commodity exports, Canada about 9 percent, and New Zealand 1 percent. (For non-energy commodities only: the shares are 6.7%, 10%, and 1.6% respectively.) Furthermore, substitution across various commodities also mitigates the market power these countries have, even within the specific market they dominate.

[32] For Australia and New Zealand, results based on other anchor currencies confirm this finding. The Canadian results using other base currencies are all insignificantly different from zero, unlike the one reported here under the IV specification.

Table III: Representative Regressions with Instrumental Variables

Dependent Variable: Log Real Exchange Rate

$$\ln(\text{Real Exchange Rate})_t = \alpha + \beta{*}t + \gamma{*}\ln(\text{Real Commodity Price})_t + \varepsilon_t$$

	Australian vs. U.S. Dollar		Canadian Dollar vs. Non-Dollar Basket		New Zealand Dollar vs. British Pound	
	OLS	GMM IV[1]: World Commodity Price[2]	OLS	GMM IV: World Commodity Price	OLS	GMM IV: World Commodity Price
Ln(Real Commodity Price)	0.81 * (0.12)	0.90 * (0.17)	-0.31 (0.21)	-0.69 * (0.23)	1.25 * (0.24)	1.97 * (0.40)
OLS: Adj. R^2 IV: 1^{st} Stage R^2	0.57	0.91	0.46	0.89	0.55	0.79
N Obs.	70		85		62	
Sample Period	1984Q1 – 2001Q2		1980Q1 – 2001Q1 [3]		1986Q1 – 2001Q2	

Note: * indicates significance at the 5% level. Newey-West heteroskedasticity and autocorrelation consistent (HAC) standard errors in parentheses.

1. Instrumental variable estimations are performed under 2SLS with GMM standard errors, using Bartlett kernel and variable Newey-West bandwidth.
2. The world commodity price index of *all* commodities is used as an instrument for the country-specific commodity price in the IV specifications. The world price index is the "non-fuel primary commodity price index" from the IMF. It consists of the US dollar prices of about 40 globally traded commodities, each weighted by their 1987-98 average world export earnings.
3. The Canadian sample here is limited to 1980Q1 to 2001Q1, the period over which world commodity price data is available.

4.2. Do Commodity Prices Better Capture Exogenous Terms of Trade Shocks Than Conventional Measures?

While previous studies have tried to incorporate terms of trade shocks into empirical exchange rate estimations of major currencies, sluggish nominal price adjustments and minimal pass-throughs typically make proper identification close to impossible. That is, in the case of sticky producer prices and perfect pass-throughs, the terms of trade and the real exchange rate will move one-to-one mechanically with no causal interpretation. The same is true when all goods are priced in local currencies, though the correlation will be of the opposite sign. When a mixture of the two pricing behavior co-exists, any sign is possible, and the dynamics are likely to be complex (see Obstfeld-Rogoff 2000). For these large commodity exporters, however, because commodity trading is mostly conducted in a few global exchange markets using U.S. dollars, world commodity price fluctuations can help us get around the identification problem to better capture the exogenous component in the variation of their terms of trade. We here consider an alternative specification: using world commodity prices as an instrument for the standard terms of trade measures. Results presented in Table IV below indicate that this idea, while theoretically sound, does not appear to have much merit empirically. From the OLS regressions, we see that (again with Canada being the exception) terms of trade indeed appear well correlated with real exchange rates. To address the endogeneity issue, country-specific world market price indices of both energy and non-energy commodities are used as instruments for terms of trade, capturing potential shocks through both import and export channels. For New Zealand, even though over half of its exports are in commodities, the low Wald statistic in the first stage regression shows that standard terms of trade measure doesn't seem to respond much to movements in the two commodity price indices.[33] For Australia, despite valid first stage regression results connecting terms of trade movements to commodity prices, the Hansen (1982) J-test rejects the over-identification restrictions, indicating that the instruments are not orthogonal to the second stage residuals, and invalidating the estimated model. We take both of these findings as support for our original specifications and our view that world commodity prices appear much better at capturing the theoretical concept of exogenous terms of trade shocks for these countries. These results also suggest that standard terms of trade measures should be used with caution in empirical exchange rate estimations, despite their conceptual importance.

[33] The coefficients on energy and non-energy commodity price index individually are not significantly different from zero either.

Table IV: Real Exchange Rates, Terms of Trade, and Commodity Prices

Dependent Variable: Log Real Exchange Rate vs. U.S. Dollar

$\ln(\text{Real Exchange Rate})_t = \alpha + \beta * t + \gamma * \ln(\text{Terms of Trade})_t + \varepsilon_t$

	Australia		Canada		New Zealand	
	OLS	GMM IV: Commodity Price [a]	OLS	GMM IV: Commodity Price	OLS	GMM IV: Commodity Price
Ln(Terms of Trade)	0.73* (0.33)	1.40 * (0.60)	-0.04 (0.20)	0.54 (0.46)	1.01 * (0.50)	3.41 (2.44)
OLS: Adj. R^2	0.16		0.58		0.23	
IV: 1st Stage Wald		p-value = 0.00		p-value = 0.00		p-value = 0.16
IV: OverID J-stats [b]		p-value = 0.01		p-value = 0.07		p-value = 0.08
N Obs.	70		118		62	
Sample Period	1984Q1 – 2001Q2		1972Q1 – 2001Q1 [3]		1986Q1 – 2001Q2	

Note: * indicates significance at the 5% level. Newey-West heteroskedasticity and autocorrelation consistent (HAC) standard errors in parentheses.

a. Instrumental variable estimations are performed under 2SLS with GMM standard errors, using Bartlett kernel and variable Newey-West bandwidth. Both country-specific energy and non-energy commodity price indices are used as instruments.

b. The J-statistics of Hansen (1982) test the null hypothesis that the GMM over-identification restrictions are satisfied/valid.

5. A Structural Interpretation of the Coefficients

Given the remarkable consistency in the estimated sign and size of the commodity price elasticity of real exchange rate, it is worth briefly considering the predictions of a simple theoretical model. Consider the following extension of the version of the Belassa-Samuelson model exposited in Obstfeld and Rogoff (1996, ch. 4). Let Home be a small economy whose agents consume three goods – nontraded goods, exports and imports – but only produce the first two. Assume that labor is perfectly mobile across industries, and that physical capital can be freely imported from abroad at real interest rate *r*, measured in importables. The production function for the exportables is

$$y_X = A_X f(k_X),$$

where y and k are output and capital per unit labor, and

$$y_N = A_N f(k_N),$$

is the analogous function for nontraded goods production. Let p_x be the world price of exportables, which is given exogenously to the small country, and p_N be the home price of nontradables, both measured in terms of importables. Then, assuming that labor mobility leads to a common wage across the two home industries, and following steps analogous to pages 205-206 in Obstfeld and Rogoff, one can derive the approximate relation:

$$\hat{p}_N = \left(\frac{\mu_{LN}}{\mu_{LX}}\right)(\hat{A}_X + \hat{p}_X) - \hat{A}_N$$

where a "hat" above a variable represents logarithmic derivatives, and μ_{LN} and μ_{LX} are the labor's income share in the nontraded and export goods sectors, respectively. Thus, the effects of a rise in the relative price of exportables is the same as a rise in traded goods productivity in the standard Belassa-Samuelson model. If $\mu_{LN} = \mu_{LX}$, a rise in the price of exportables leads to a proportional rise in the price of nontraded goods. The impact on the real exchange rate depends, of course, on the utility function. Assume a simple logarithmic (unit-elastic) utility function:

$$U = C_N^{\alpha} C_I^{\beta} C_X^{(1-\alpha-\beta)}$$

Normalizing the price of importables to one, the consumption-based consumer price index is then given by

$$p_N^{\alpha} p_X^{(1-\alpha-\beta)}$$

Therefore, as $\hat{p}_N$ moves proportionately in response to $\hat{p}_X$, the effect of an export price shock on the utility-based real CPI is then given by

$$\hat{p}_X^{\ (1-\beta)}$$

Assuming that importables account for *25*% of consumption, then the elasticity of the CPI with respect to a unit change in the price of exportables would then be *0.75*, which is broadly consistent with our estimated coefficients. (If $\mu_{LN} > \mu_{LX}$ – it is standard to assume that nontraded goods production is labor intensive -- one gets a larger effect).

What if the price of nontraded goods is sticky? Then a simple model of optimal monetary policy would predict that the exchange rate should be adjusted one for one with changes in the world price of exportables, in order to accommodate the requisite rise in the relative price of nontradable goods. (This assumes that export prices are flexible with complete pass-through, as otherwise a larger change in the exchange rate would be needed.) Of course, if the central bank is mechanically trying to stabilize CPI inflation, and if its rule does not allow any offset for export price shocks, then the authorities would not allow the nominal exchange rate to move by the amount required to mimic the flexible-price equilibrium, but instead only by a smaller amount.

We have only offered one model, but many others can give parallel results. For example, the classic model of Dornbusch (1976) would also prescribe a one-for-one movement of the exchange rate in response to terms of trade shocks, in order to mimic the real allocation of the flexible price equilibrium. Whereas we have no illusions that the simple model presented here fully describes the data, it still provides a useful benchmark for assessing the estimated coefficients.

6. Empirical Exchange Rate Puzzles

While canonical exchange rate models such as Dornbusch's (1976) overshooting model seemed to broadly fit the facts for the 1970s and the early 1980s, as inflation gradually stabilized in major OECD countries over the ensuing period, it became clear that monetary instability *alone* could not possibly explain the persistent exchange rate volatility that remains even to this date. The failure of standard monetary models further resonates in their inability to reconcile the extremely slow rate at which deviations from PPP seem to die out, with the enormous short-term volatility observed in real exchange rates. As exposited in Rogoff (1996), conventional shocks to the real economy such as taste or technology shocks, while capable of generating slow adjustment, are simply not volatile enough to account for the short-term variation in the exchange rates. Models based on monetary or financial shocks may explain this short-term volatility, but the long half-lives of shocks observed in the data are incompatible with the concept of long-run monetary neutrality under these models. Hence, a potential solution to this PPP puzzle may lie in identifying a shock that is both sufficiently volatile and persistent.

The success of our univariate regressions suggests that commodity prices may indeed be this missing shock. Would the Dornbusch-type monetary variables work better in these rare country cases where certain real shocks can be substantially controlled for? An affirmative answer would require sufficiently removing the persistence in real exchange rate shocks, so as to allow monetary variables to account for the remaining variations.[34] By examining the degree of persistence in real exchange rate shocks, we show in this section that commodity prices are no Deus ex Machina; that is, although they are found to be a strong and consistent explanatory variable in exchange rate equations, their introduction do not otherwise resuscitate the monetary approach to exchange rate, at least from an empirical perspective.[35]

[34] Here we ignore the possibility of non-linear adjustment to PPP but focus on linear models.
[35] Indeed, incorporating commodity prices into standard monetary-type regressions only corroborates further the "fickleness" of standard models documented in the literature, and provides little support for a commodity price augmented Dornbusch model. This section looks at the reason why monetary fundamentals in the Dornbusch model may be inappropriate in explaining the remaining variation in our augmented exchange rate equations.

6.1. "The Nagging Persistence"

To examine the degree of persistence in real exchange rates, we assume real exchange rate shocks to follow an AR(1) process and focus on the magnitude of the autoregression coefficients.[36] We have already seen in Table 1 that the AR roots are very large in the commodity price equations; Table V below presents a more systematic analysis. The AR(1) columns in Table V demonstrate the persistence side of the standard PPP puzzle (sans commodity prices) for the three commodity currencies. The estimated autocorrelations in the residuals appear broadly similar to what we see in the PPP literature, indicating half-lives much longer than what monetary factors can explain. We note that OLS estimates of the AR coefficients are well known to have substantial bias, especially when the autocorrelation is close to unity and the sample size is small.[37] Work by Andrews (1993) and Fair (1996), among others, examine this bias extensively and propose various methods of correction.[38] Furthermore, the direction of the bias has also demonstrated to be downward towards zero.[39] So, while our reported point estimates for the AR roots are biased, the underestimated degree of persistence is nevertheless high enough to provide meaningful insight to the nature of the "puzzle" at hand.[40]

Having established the PPP puzzle in our specific currencies, we then control for the effects of commodity prices – removing exchange rate variations due to commodity price shocks – to see if the persistence lessens significantly. As evident from Table V, even after controlling for commodity price shocks and instrumenting for potential endogeneity, exchange rate residuals still exhibit a generally similar degree of persistence.[41] As the implied half-lives from these

[36] See Froot and Rogoff (1995) and Rogoff (1996) for discussions of previous literature using this specification and other variants. There are certainly alternative methods for capturing exchange rate persistence.
[37] See, for example, Mark (2001) and Murray and Papell (2002) for a discussion of these biases in relation to the PPP half-life literature.
[38] These papers propose using variants of median-unbiased estimators for autoregressive coefficients.
[39] As discussed in Murray and Papell (2002), the LS bias is always downward in the AR(1) model. For higher-order AR specifications, the high degree of persistence observed in real exchange rates should also be sufficient to ensure downward biases. (See Stine and Shaman (1989) for details on potential "bias direction reversal".)
[40] Of course, the precision of these point estimates is another thorny issue. We recognize that the confidence intervals, which can be constructed via various bootstrap methods, are likely to be extremely wide; however, this is the limitation of analysis based on small sample sizes like the ones we have.
[41] We also examined the adjustment dynamics of real exchange rates through impulse response analysis, allowing for possible higher order autocorrelation structures, hence potential non-monotonic responses to shocks (See Cheung and Lai (2000) or Murray and Papell (2002)). The dynamic response patterns show that incorporating higher order AR terms do not significantly alter the persistence of shocks obtained under the AR(1) specifications. We note again that these persistence estimates are extremely imprecise, given our small sample size.

coefficients are far longer than one can justify if the main source of the remaining shocks is monetary, it is no surprise that, as mentioned in footnote 35, we saw little empirical support for commodity price augmented Dornbusch-type equations.

Table V: Persistence in the Real Exchange Rates [1]

Dependent Variable: Log of Real Exchange Rate, vs. U.S. Dollar

	Australia			Canada			New Zealand		
	AR(1) [2]	AR(1) + Comm Price [3]	AR(1) + IV [4]	AR(1)	AR(1) + Comm Price	AR(1) + IV	AR(1)	AR(1) + Comm Price	AR(1) + IV
Ln(Real Commodity Prices)		0.54 * (0.14)	0.97 * (0.35)		0.04 (0.08)	-0.27 (0.16)		0.51 * (0.21)	0.72 * (0.35)
AR(1) root	0.94 * (0.04)	0.88 * (0.05)	0.84 * (0.08)	0.97 * (0.03)	0.96 * (0.03)	0.97 * (0.03)	0.92 * (0.04)	0.95 * (0.05)	0.95 * (0.06)
Durbin-Watson stat	2.03	2.15	2.06	1.98	2.01	1.67	2.03	1.99	1.92
Adj. R^2 / 1^{st} Stage R^2	0.84	0.86	0.94	0.95	0.95	0.89	0.90	0.90	0.92
	70			86			62		
Sample Period	1984Q1 – 2001Q2			1980Q1 - 2001Q2			1986Q1 - 2001Q2		

Note: * indicates significance at the 5% level.

1. The AR root estimates in this table are downward biased (towards zero); see text for discussion.
2. ln(Real Exchange Rate)$_t$ = $\alpha + \beta * t + \varepsilon_t$, where ε_t follows an AR(1).
3. ln(Real Exchange Rate)$_t$ = $\alpha + \beta * t + \gamma *$ln(Real Commodity Price)$_t$ + ε_t, where ε_t follows an AR(1).
4. ln(Real Exchange Rate)$_t$ = $\alpha + \beta * t + \gamma *$ln(Real Commodity Price)$_t$ + $\varepsilon_{t,}$ where ε_t follows an AR(1) and ln(World Commodity Price Index) is used as an IV.

6.2. Other Shocks: The Balassa-Samuelson Relative Productivity Differences

As the commodity price-exchange rate connection appears more a “Down Under” phenomenon, here we analyze the Australia and New Zealand real exchange rate further by identifying and controlling for an additional source of real shocks. As discussed in Section 5, the Balassa-Samuelson model suggests that country differences in traded and non-tradable sector productivity shocks may affect real exchange rate movements through their impact on relative wages. Figure 7 plots Australian and New Zealand real exchange rates with the ratios of the home country traded versus non-traded sector productivity to that of the United States.[42] In contrast to the real interest rate and output differentials series presented in Figure 1a and 3a, we see much more obvious visual correlations between relative productivity differences and the real exchange rates. Indeed, results in Table VI below show coefficient estimates roughly consistent with predictions of the Balassa-Samuelson framework presented in Section 5. However, the productivity measures appear to be a less robust explanatory variable than commodity prices. If the Newey-West procedure is replaced by the parametric AR(1) specification, relative productivity no longer shows up as significant, while commodity prices remain resilient. More importantly, we note that the AR root coefficient estimates in the "commodity price cum relative productivity-augmented equations" remain high as before.[43] Hence, we find the PPP puzzle to be like the Russian dolls, in that after controlling for *two* promising real shocks – peeling away two layers of the original PPP puzzle – we are still faced with the identical, despite smaller, PPP puzzle.

[42] We were unable to obtain consistent productivity measures across countries, but they are consistent across sectors within a country. This is not ideal, but as we look at differences in within-country productivity ratios, we think the inconsistency is not a serious problem. See the Data Appendix for the details on these variables.
[43] Although these AR(1) coefficients are measured with imprecision, the magnitude is nevertheless very large for estimators that are downward biased.

Table VI: Productivity Differentials and Real Exchange Rates [1]

Dependent Variable: Log of Real Exchange Rate, vs. U.S. Dollar

$\ln(\text{Real Exchange Rate})_t = \alpha + \beta{*}t + \gamma_1{*}\ln(\text{Real Commodity Price})_t + \gamma_2{*}$ Diffierential of ln(Tradable/ Non-Tradable Productivity) vs. the U.S.$)_t + \varepsilon_t$

	Australia				New Zealand			
	OLS + Newey-West	OLS+ AR(1)	IV: World Comm Price	IV + AR(1)	OLS + Newey-West	OLS+ AR(1)	GMM IV: World Comm Price	IV + AR(1)
Ln(Real Commodity Prices)	0.75 * (0.10)	0.56 * (0.20)	0.86 * (0.15)	1.08 * (0.34)	1.09 * (0.29)	0.53 * (0.20)	2.16 * (0.56)	0.83 * (0.30)
Tradable-Non-Tradable Prod. Diff	0.87 * (0.29)	-0.03 (0.15)	0.83 * (0.34)	0.10 (0.15)	0.90 * (0.38)	0.14 (0.17)	0.87 * (0.43)	0.16 (0.17)
AR(1) root [1]		0.89 * (0.05)		0.83 * (0.08)		0.95 * (0.05)		0.96 * (0.06)
Durbin-Watson stat	0.66	1.97	0.63		0.27	1.99	0.29	
Adj. R^2 / 1st Stage R^2	0.66	0.86	0.95	0.95	0.44	0.90	0.92	0.92
	67				62			
Sample Period	1984Q4 - 2001Q2				1986Q1 - 2001Q2			

Note: * indicates significance at the 5% level.

5. The AR root estimates in this table are downward biased (towards zero); see text for discussion.

1. $\ln(\text{Real Exchange Rate})_t = \alpha + \beta{*}t + \varepsilon_t$, where ε_t follows an AR(1).
2. $\ln(\text{Real Exchange Rate})_t = \alpha + \beta{*}t + \gamma{*}\ln(\text{Real Commodity Price})_t + \varepsilon_t$, where ε_t follows an AR(1).
3. $\ln(\text{Real Exchange Rate})_t = \alpha + \beta{*}t + \gamma{*}\ln(\text{Real Commodity Price})_t + \varepsilon_t$, where ε_t follows an AR(1) and ln(World Commodity Price Index) is used as an IV.

7. Conclusion

In a literature largely populated by negative findings and empirical puzzles, this paper identifies a source of exogenous shocks and explores its contribution to time series exchange rate behavior, and more broadly, with standard exchange rate models. The world prices of commodity exports, measured in real U.S. dollars, do appear to have a strong and stable influence on the real exchange rates of New Zealand and Australia. For Canada, the relationship is somewhat less robust, especially to detrending. Thus, despite the fact that these countries had open capital markets and free floating exchange rates over the sample period, one can identify an important real explanatory variable. Moreover, the quantitative size of the coefficient is broadly consistent with the predictions of standard theoretical models of optimal monetary policy.

Although Australia, Canada and New Zealand are fairly unique among OECD countries, commodity price shocks (both export and import) have long been recognized as of great importance to many developing countries that rely heavily on primary commodity production. The experience of Australia, Canada, and New Zealand are of particular relevance as many of these developing countries liberate capital markets and move towards floating exchange rate systems. While this paper mainly covers the empirical links, understanding exchange rate responses to world commodity price shocks can provide important information for a broad range of policy issues, including especially the conduct of monetary policy and inflation control.

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Appendix:

Table A.1

Commodity Price Elasticities of Real Exchange Rates

Dependent Variable: Log of Real Exchange Rate, relative to Different Anchor Currencies

$$\ln(\text{Real Exchange Rate})_t = \alpha + \beta * \ln(\text{Real Commodity Price})_t + \varepsilon_t$$

	Australia			Canada			New Zealand		
National Currency	vs. U.S. Dollar	vs. British Pound	vs. Non-Dollar Basket [1]	vs. U.S. Dollar	vs. British Pound	vs. Non-Dollar Basket	vs. U.S. Dollar	vs. British Pound	vs. Non-Dollar Basket
Ln(Real Commodity Prices) [2]	0.40 * (0.08)	0.51 * (0.09)	0.36 * (0.06)	0.40 * (0.07)	0.50 * (0.14)	0.36 * (0.09)	0.53 * (0.17)	0.61 * (0.10)	0.41 * (0.10)
Adj. R^2	0.39	0.55	0.51	0.56	0.34	0.34	0.30	0.45	0.36
N Obs.	70			114			62		
Sample Period	1984Q1 – 2001Q2			1973Q1 - 2001Q2			1986Q1 - 2001Q2		

Note: * indicates significance at the 5% level. Newey-West heteroskedasticity and autocorrelation consistent (HAC) standard errors in parentheses.

1. “Non-dollar Basket” is a U.S. trade-weighted average of over 30 currencies, excluding the U.S. dollar, of major U.S. trading partners. It is based on the broad real index from the Federal Reserve.
2. Real commodity price index is the base country production-weighted average of world commodity prices in U.S. dollars, deflated by the U.S. CPI. See the data appendix for country-specific production weights.

Table A.2

Parameter Stability Tests using Linear Filter

Dependent Variable: Log of Real Exchange Rate, vs. U.S. Dollar

Detrended ln(Real Exchange Rate)$_t$ = $\alpha + d_t + (\beta + \gamma * d_t)$* Detrended ln(Real Commodity Price)$_t + \varepsilon_t$,

where $d_t = 1$ if $t \geq$ Breakpoint; $d_t = 0$ otherwise

Linearly detrended variables	Australia		Canada		New Zealand	
	OLS + Hansen Test	OLS + Break Dummy	OLS + Hansen Test	OLS + Break Dummy	OLS + Hansen Test	OLS + Break Dummy
Ln(Real Commodity Prices): β	0.81* (0.12)	0.74 * (0.17)	0.21 (0.15)	0.38 * (0.16)	1.10 (0.32)	2.13 * (0.38)
Dummy* Ln(Real Commodity Prices): γ		0.18 (0.26)		-0.60 * (0.29)		-1.14 * (0.50)
Breakpoint tested [1]	NA	1993Q1	NA	1991Q1	NA	1988Q1
Hansen test [2]	0.37		0.49 *		0.39	
Adj. R^2	0.53	0.52	0.05	0.14	0.27	0.37
N Obs.	70		114		62	
Sample Period	1984Q1 – 2001Q2		1973Q1 - 2001Q2		1986Q1 - 2001Q2	

Note: * indicates significance at the 5% level. Newey-West heteroskedasticity and autocorrelation consistent (HAC) standard errors in parentheses.
1. The breakpoint is the starting year for the use of formal inflation targets in the country.
2. The 5% asymptotic critical value for the Hansen individual parameter test is 0.47 (see Hansen 1992, Table 1).

Table A.3

N-Period Ahead Out-of-Sample Forecasts of Real Exchange Rate Levels

Root Mean Squared Errors Ratios of Commodity Price Augmented Model vs. Random Walk*

Forecast Horizons	Australian Dollar vs.				Canadian Dollar vs.				New Zealand Dollar vs.			
	US$	British Pound	Yen	Non-$ Basket	US$	British Pound	Yen	Non-$ Basket	US$	British Pound	Yen	Non-$ Basket
1-Quarter												
1996:1 - 2001:2	1.01	0.97	1.04	0.99	1.02	1.01	1.01	1.02	1.00	1.03	1.08	1.02
1996:1 - 1999:4	0.97	0.99	1.03	1.13	1.03	1.01	0.96	1.01	0.96	1.07	1.10	1.15
1993:1 - 2001:2	1.00	1.03	1.01	0.96	1.04	1.01	0.98	1.02	1.09	1.22	1.05	1.11
4-Quarters												
1996:1 - 2001:2	0.82	0.82	1.12	1.00	0.99	1.07	0.90	0.90	1.02	1.14	1.24	1.12
1996:1 - 1999:4	0.78	0.82	1.15	1.13	1.01	1.13	0.88	0.87	0.84	1.16	1.20	1.22
1993:1 - 2001:2	0.80	1.07	1.00	0.94	1.12	0.99	0.97	1.19	1.13	1.28	1.24	1.24
8-Quarters												
1996:1 - 2001:2	0.45	0.86	0.73	0.85	1.31	1.39	0.72	1.43	0.88	0.92	1.15	1.00

* Root mean squared forecast error from the following rolling regression is compared to that of the Random Walk:

$$\ln(\text{Real Exchange Rate})_t = \alpha + \rho^* \ln(\text{Real Exchange Rate})_{t-1} + \beta^* \ln(\text{Real Energy Commodity Price})_t + \gamma^* \ln(\text{Real Non-Energy Commodity Price})_t + \varepsilon_t.$$

** The shaded numbers show where the forecast model performs better than the RW model, though we are not assessing the significance of the difference here. See text for further discussions.

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. Real exchange rates are nominal rates adjusted by the CPI ratios. To construct the real rates against the non-dollar basket, we use the broad (real) index published by the Federal Reserve to adjust the country real rates against the U.S. dollar. The broad index measures the foreign exchange value of the U.S. dollar against the currencies of a large group of U.S. trading partner.

CPI and Real Output:

- Quarterly consumer prices and GDP volume (1995 = 100) are taken from the IFS. Inflation rates are calculated as annualized quarterly changes of the CPI.

Short Term Interest Rates:

- We use three-month Treasury bill rates as a measure of short-term interest rate. The sources are IFS and Global Financial Database. The real rates are the nominal rates adjusted for expected inflation, which we proxy with lagged inflation rate, calculated as the annualized quarterly change of the CPI from the previous quarter.

Terms of Trade:

- Country-specific export and import price indices are provided by Bank of Canada, Reserve Bank of Australia, and Reserve Bank of New Zealand.

Relative Productivity of Traded to Non-Traded Sectors:

- Australia: quarterly labor productivity for the traded and non-traded parts of the market sector economy are measured as real output per hours worked. The market sector makes up about two-thirds of the overall Australian economy. Traded vs. non-traded are determined on the basis of export and/or import intensities of the industries. The following sectors are classified as traded: agriculture, forestry etc; mining; manufacturing (except wood and paper, printing/publishing and non-metallic minerals); air transport, and water transport. Non-traded sectors are: wood and paper products, printing and publishing, non-metallic minerals, utilities, construction, wholesale trade, retail trade, accommodation etc, road transport, rail and pipelines, transport services and storage, communications, finance and insurance, cultural and recreational services.

- New Zealand: productivity is defined as seasonal adjusted GDP relative to the number of people employed (Household Labor Force Survey). Tradable sectors include: agriculture, hunting, fishing & forestry; manufacturing; and mining and quarrying. Non-tradable productivity covers the following sectors of the economy: building and construction; business and financial services; community, social and personal services; electricity, gas and water; transport, storage and communication; wholesale and retail trade; and others.

- The U.S.: The productivity measure is constructed using quarterly NIPA real GDP and BLS worker-hours. Goods-producing sector is treated as traded, and service-producing sector non-traded.

Commodity Prices:

- The country specific commodity export price index is constructed by geometrically weighting the world market prices in U.S. dollar of each country's major non-energy commodity exports. The weights, taken from Djoudad, Murray, Chan, and Daw (2001), represent the average production value of the commodity over the 1982-90 period and are listed in Table A. 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 price index of all non-energy commodities is the "non-fuel primary commodity price index" of the IMF. It comprises the U.S. dollar prices of about 40 globally traded commodities, each weighted by their 1987-98 average world export earnings.
- The world market price of individual commodities is taken from various sources (see Table A). They are quarterly average spot or cash prices in U.S. 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".

Table A. Composition of Non-Energy Commodity Price Index

World Market Price in U.S. Dollar

Australia			Canada			New Zealand		
1983Q1 – 2001Q2			1972Q1 - 2001Q2			1986Q1 - 2001Q2		
Product	Wt.	Source	Product	Wt.	Source	Product	Wt.	Source
Aluminum	9.1%	IMF	Aluminum	4.8%	BOC	Aluminum	8.3%	ANZ
Beef	9.2%	IMF	Beef	9.8%	GFD	Apples	3.1%	ANZ
Copper	3.2%	BOC	Canola	2.1%	BOC	Beef	9.4%	ANZ
Cotton	3.4%	IMF	Copper	4.7%	BOC	Butter	6.5%	ANZ
Gold	19.9%	IMF	Corn	1.3%	BOC	Casein	6.7%	ANZ
Iron Ore	10.9%	IMF	Gold	4.5%	GFD	Cheese	8.3%	ANZ
Lead	1.3%	IMF	Hogs	5.1%	GFD	Fish	6.7%	ANZ
Nickel	2.6%	BOC	Lumber	14.4%	IMF	Kiwi	3.7%	ANZ
Rice	0.8%	IMF	Newsprint	13.4%	IMF	Lamb	12.5%	ANZ
Sugar	5.9%	GFD	Nickel	3.9%	BOC	Logs	3.5%	ANZ
Wheat	13.5%	BOC	Potash	2.1%	IMF	Pulp	3.1%	ANZ
Wool	18.3%	ANZ + IMF	Pulp	19.7%	IMF	Sawn Timber	4.6%	ANZ
Zinc	1.8%	BOC	Silver	0.9%	GFD	Skim MP	3.7%	ANZ
			Wheat	8.9%	BOC	Skins	1.6%	ANZ
			Zinc	4.4%	BOC	Wholemeal MP	10.6%	ANZ
						Wool	7.7%	ANZ

Note: BOC (Bank of Canada); ANZ (Australia-New Zealand Bank); GFD (Global Financial Database)

Fig. 1a: US - Australian Real Exchange Rate, Real interest Differential, Real Output Differential, 1984q1 to 2001q2

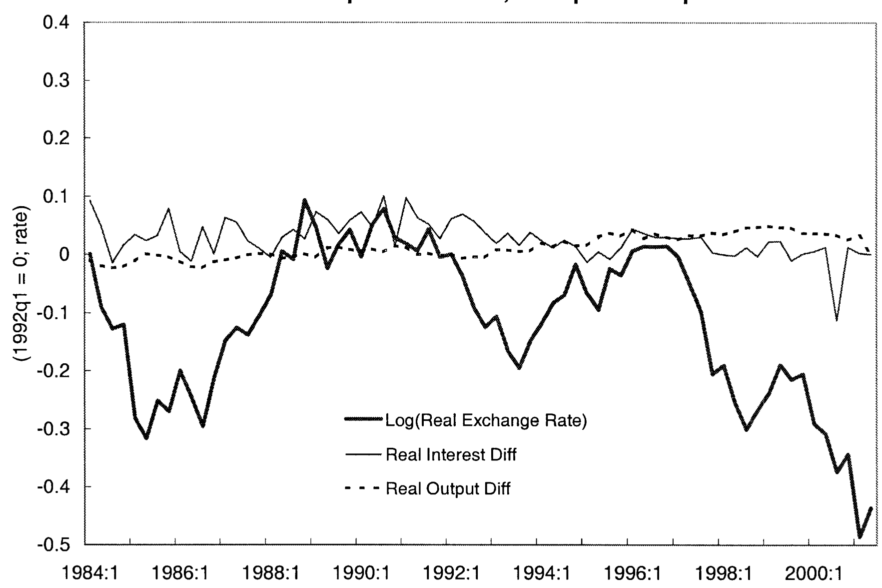

Fig. 1b: US - Australian Real Exchange Rate, Real Commodity Price, 1984q1 to 2001q2

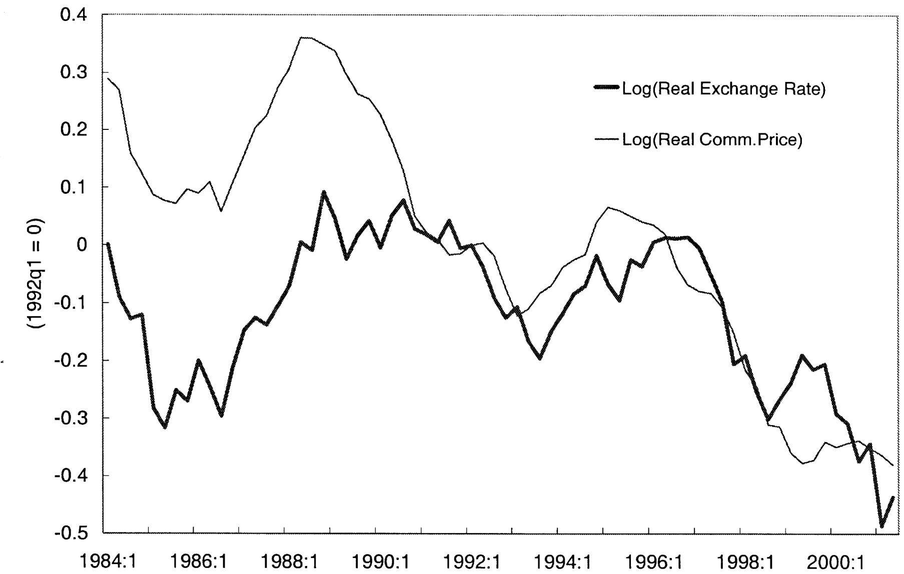

Fig. 1c: Non-Dollar Basket - Australian Real Exchange Rate, Real Commodity Price, 1984q1 to 2001q2

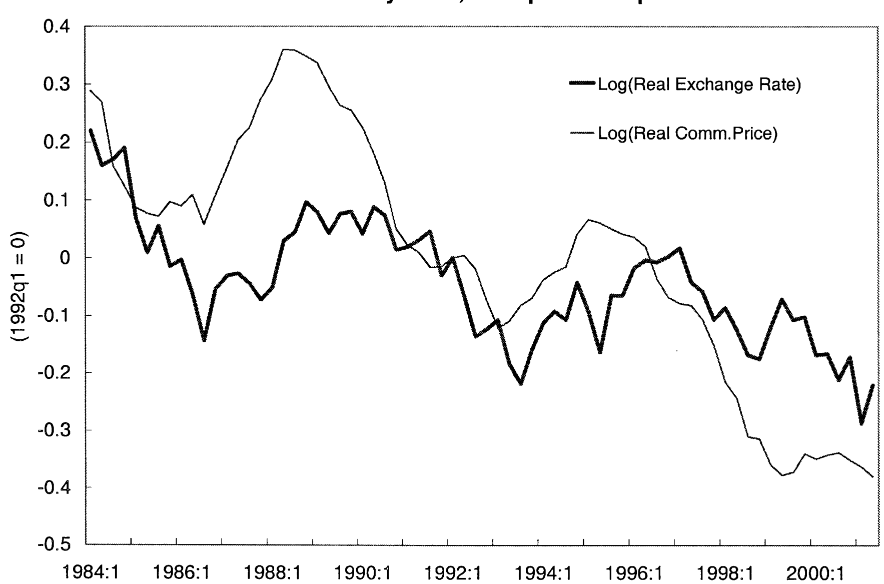

Fig. 1d: UK - Australian Real Exchange Rate, Real Commodity Price, 1984q1 to 2001q2

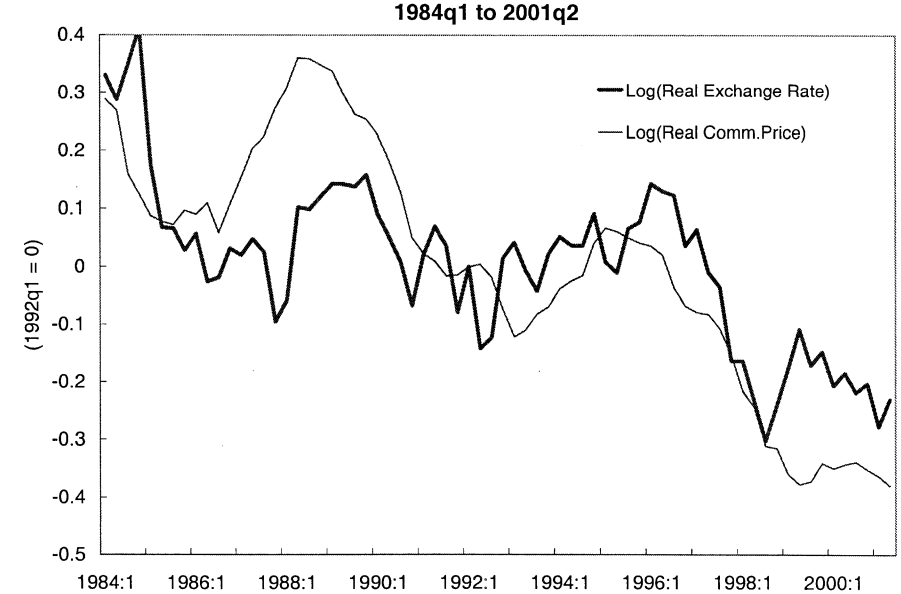

Fig. 2a: US - Canadian Real Exchange Rate, Real Interest Differential, Real Output Differential, 1973q1 to 2001q2

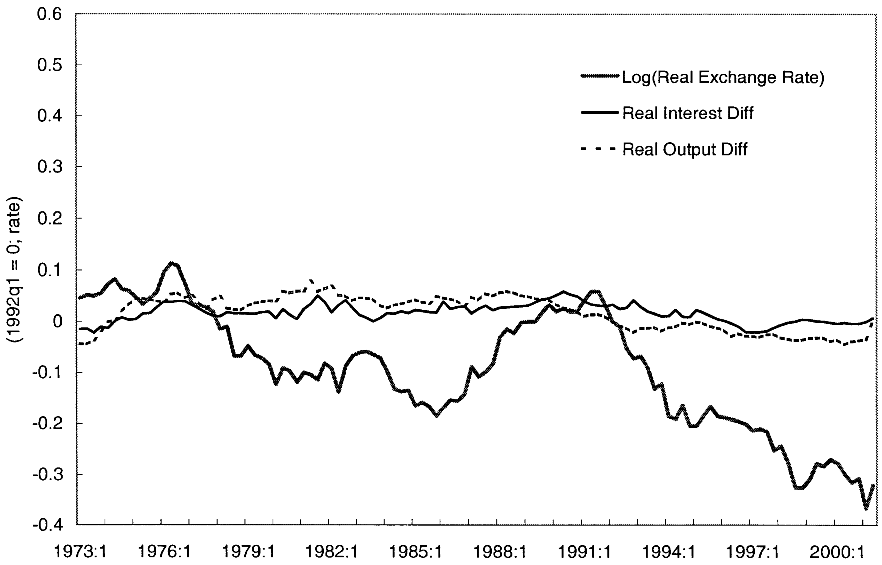

Fig. 2b: US - Canadian Real Exchange Rate, Real Commodity Price 1973q1 to 2001q2

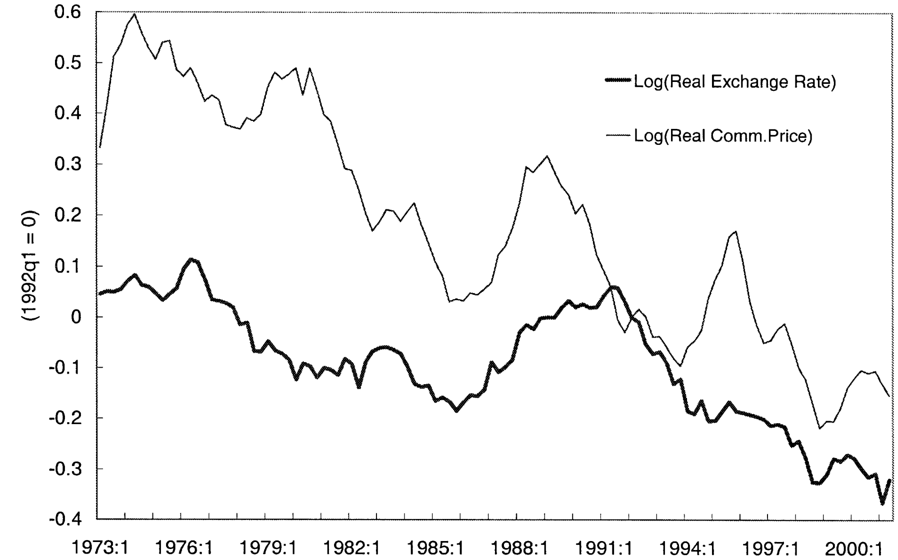

Fig. 2c: Non-Dollar Basket - Canadian Real Exchange Rate, Real Commodity Price, 1973q1 to 2001q2

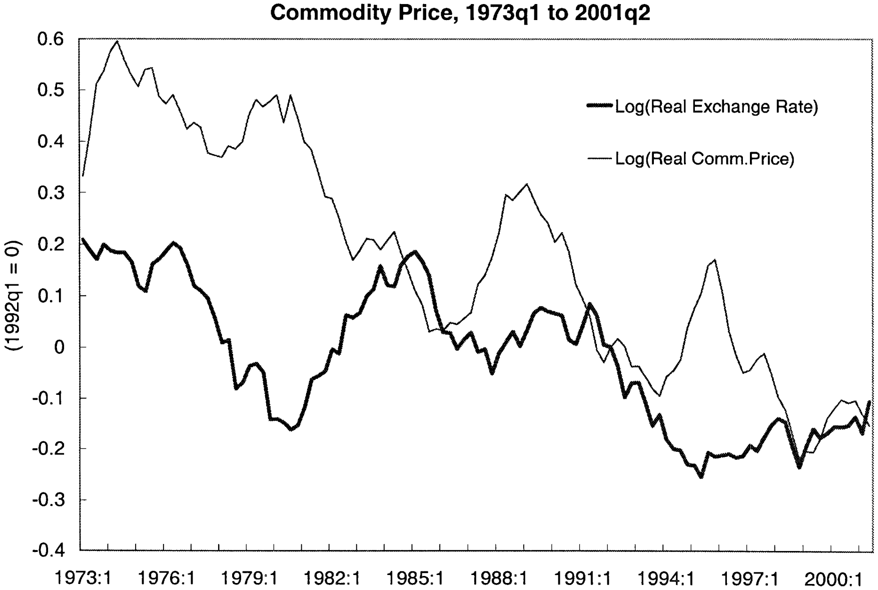

Fig. 2d: UK - Canadian Real Exchange Rate, Real Commodity Price, 1973q1 to 2001q2

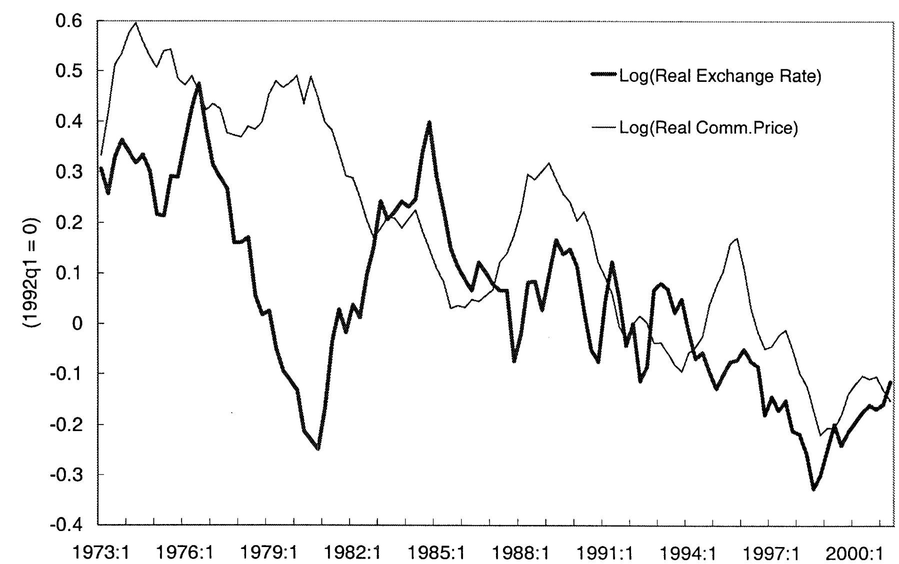

Fig. 3a: US - New Zealand Real Exchange Rate, Real Interest Differential, Real Output Differential, 1986q1 to 2001q2

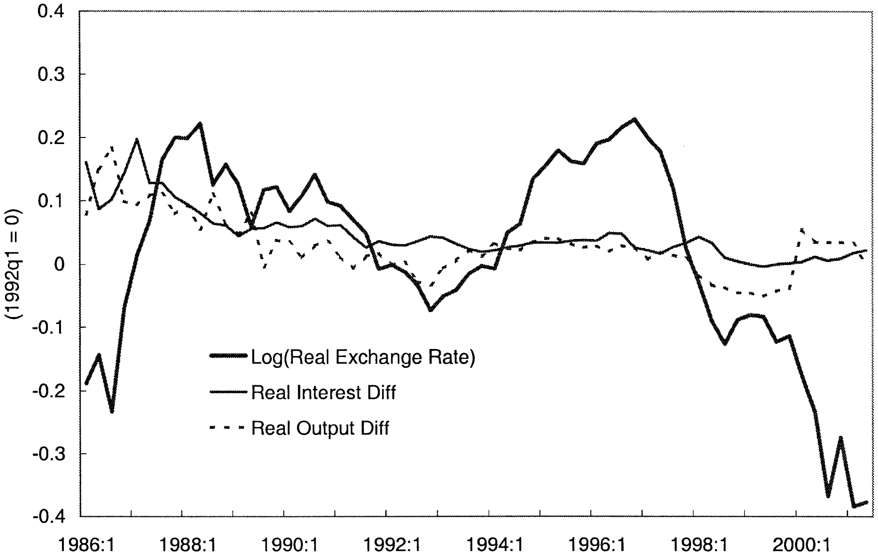

Fig. 3b: US - New Zealand Real Exchange Rate, Real Commodity Price, 1986q1 to 2001q2

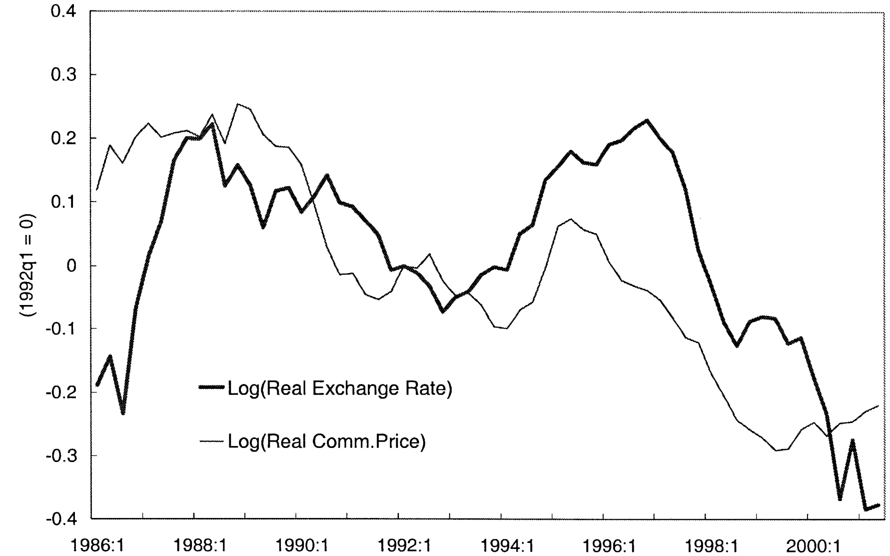

Fig. 3c: Non-Dollar Basket - New Zealand Real Exchange Rate, Real Commodity Price, 1986q1 to 2001q2

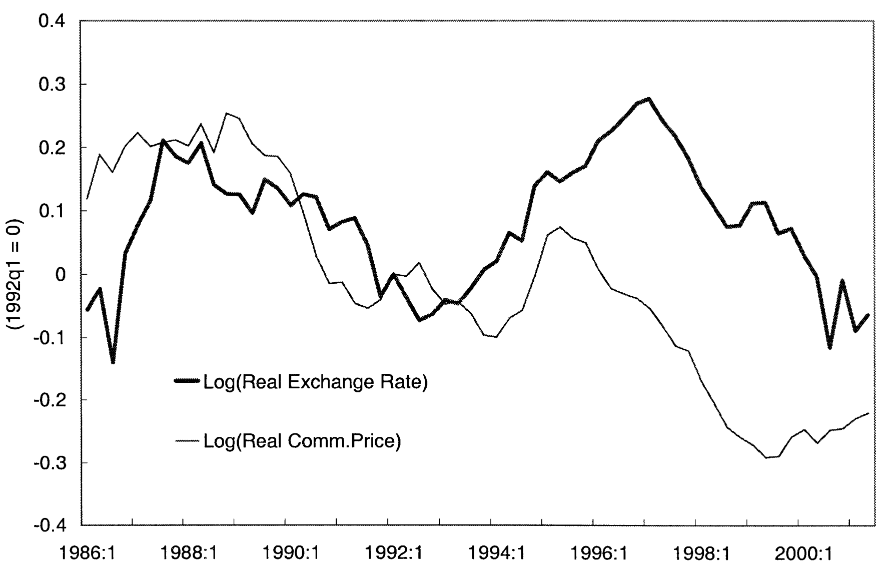

Fig. 3d: UK - New Zealand Real Exchange Rate, Real Commodity Price, 1986q1 to 2001q2

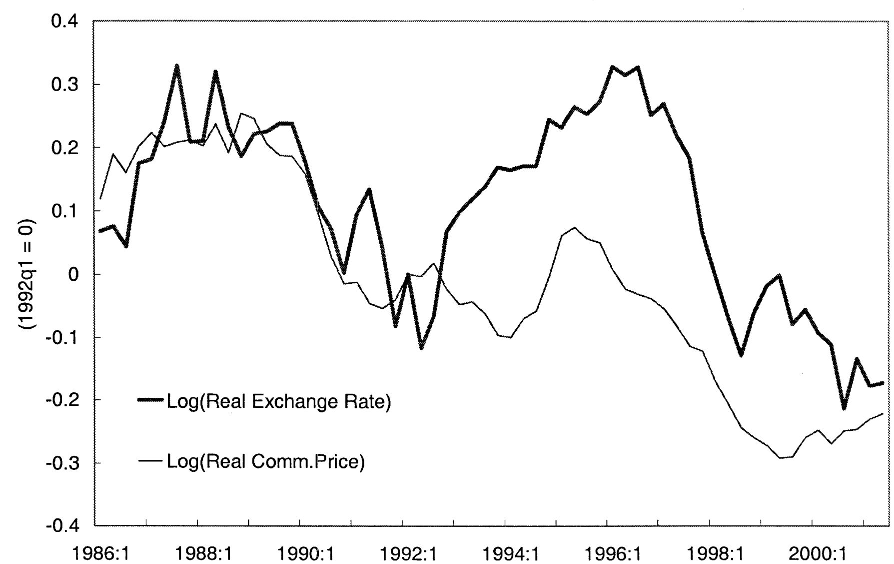

Fig. 4a: Detrended US - Australia Real Exchange Rate, Real Commodity Price, 1984q1 to 2001q2

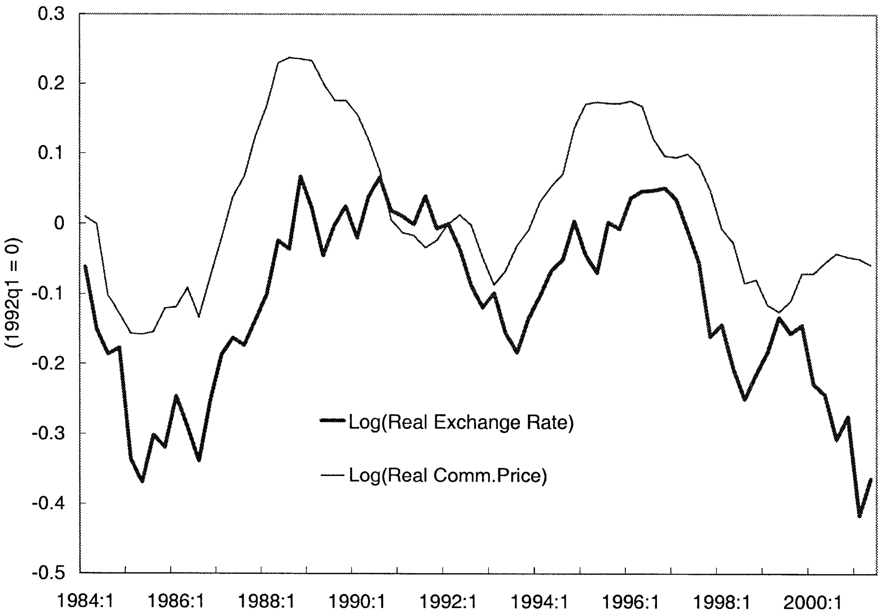

Fig. 4b: Detrended Non-Dollar Basket - Australian Real Exchange Rate, Real Commodity Price, 1984q1 to 2001q2

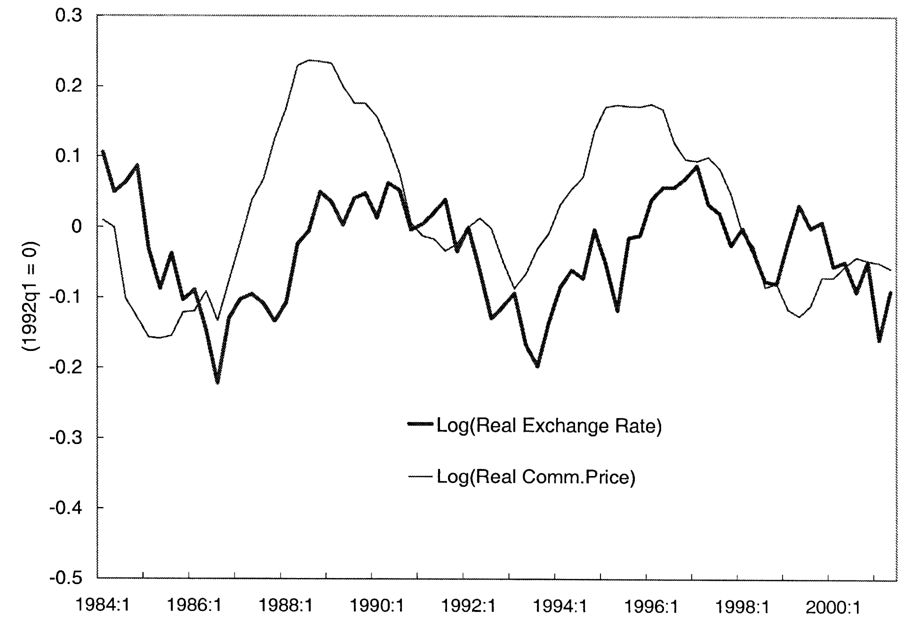

Fig. 4c: Detrended UK - Australia Real Exchange Rate, Real Commodity Price, 1984q1 to 2001q2

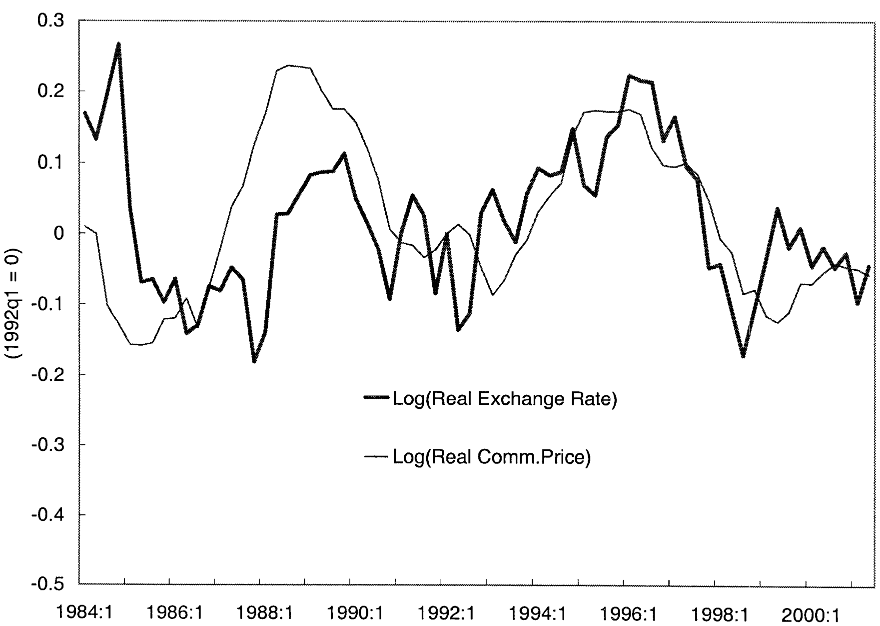

Fig. 5a: Detrended US - Canadian Real Exchange Rate, Real Commodity Price, 1973q1 to 2001q2

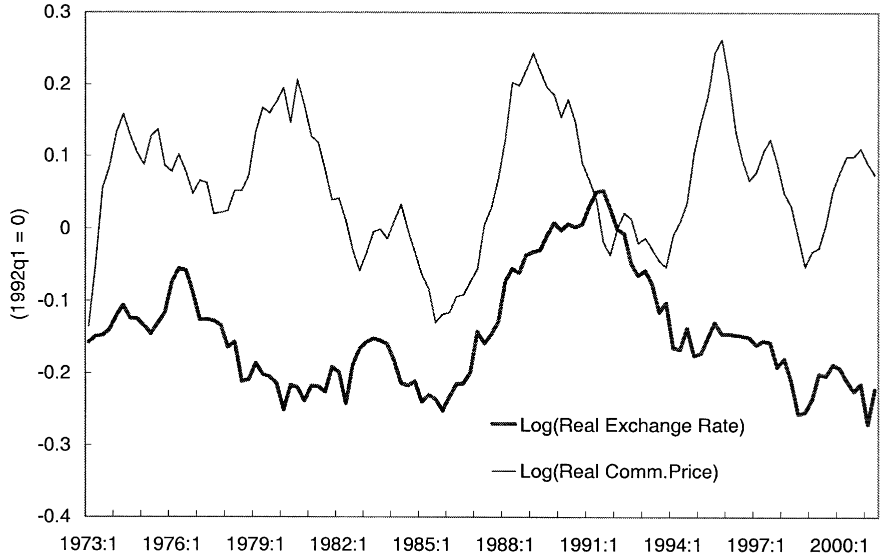

Fig. 5b: Detrended Non-Dollar Basket - Canadian Real Exchange Rate, Real Commodity Price, 1973q1 to 2001q2

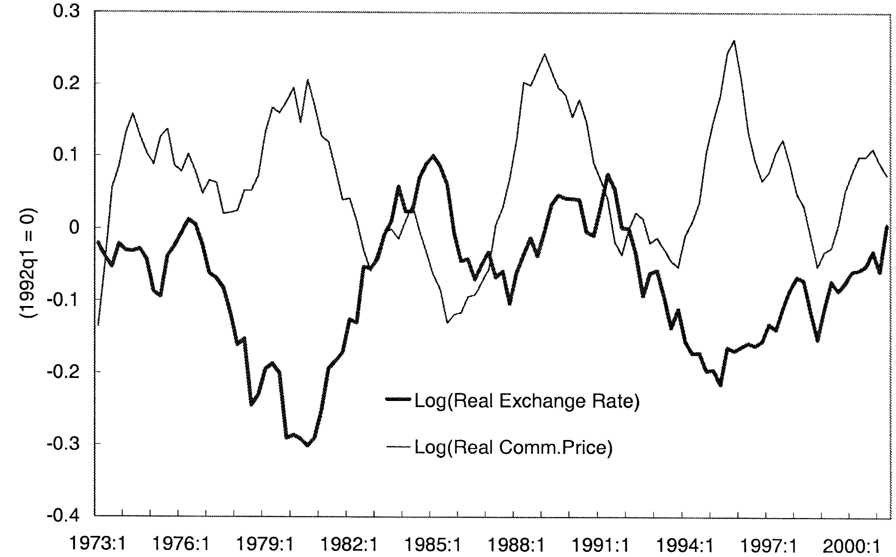

Fig. 5c: Detrended UK - Canadian Real Exchange Rate, Real Commodity Price, 1973q1 to 2001q2

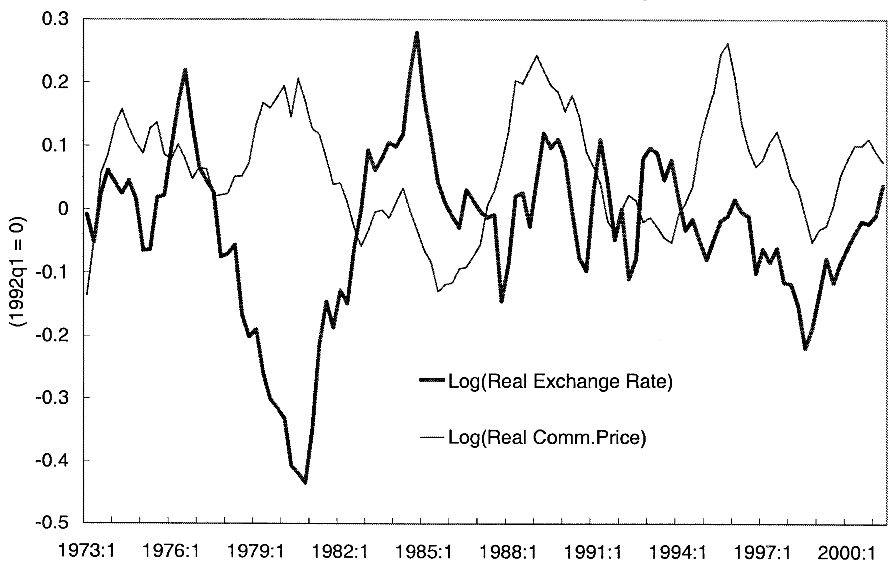

Fig. 6a: Detrended US - New Zealand Real Exchange Rate, Real Commodity Price, 1986q1 to 2001q2

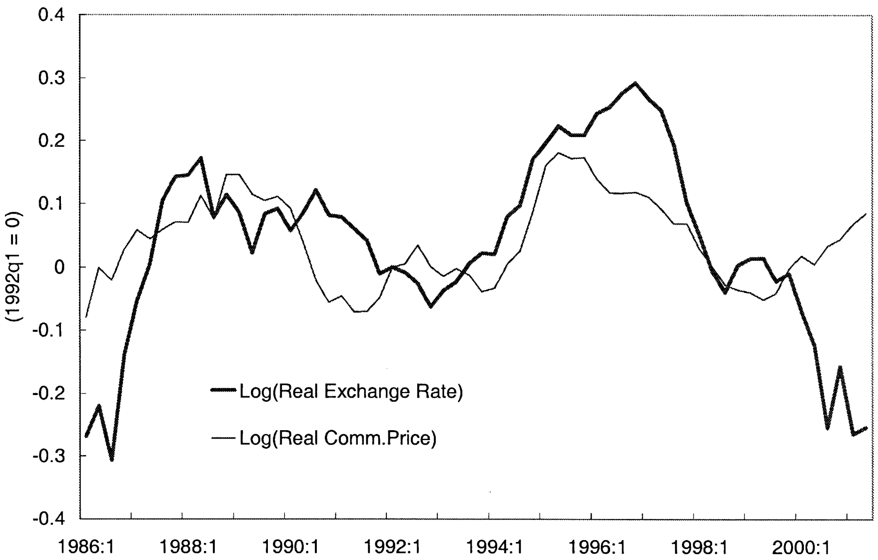

Fig. 6b: Detrended Non-Dollar Basket - NZL Real Exchange Rate, Real Commodity Price, 1986q1 to 2001q2

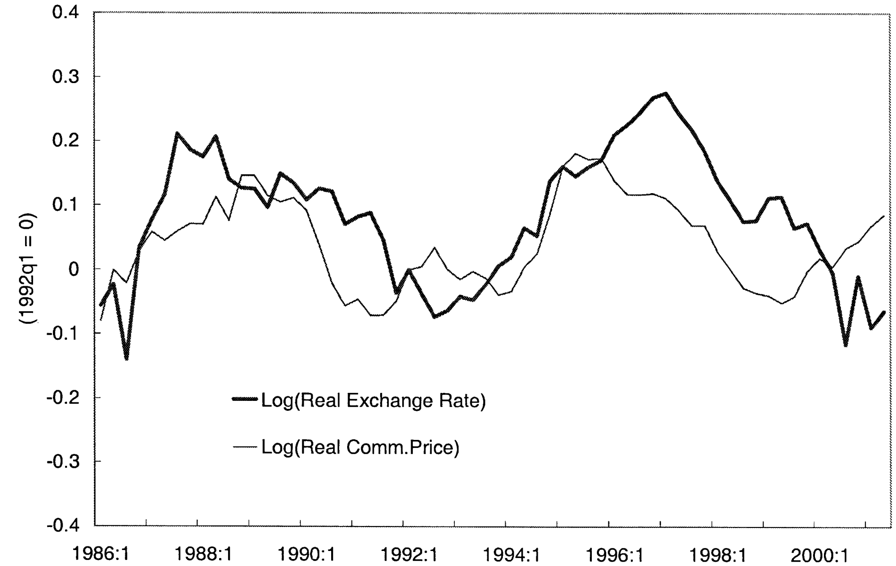

Fig. 6c: Detrended UK - New Zealand Real Exchange Rate, Real Commodity Price, 1986q1 to 2001q2

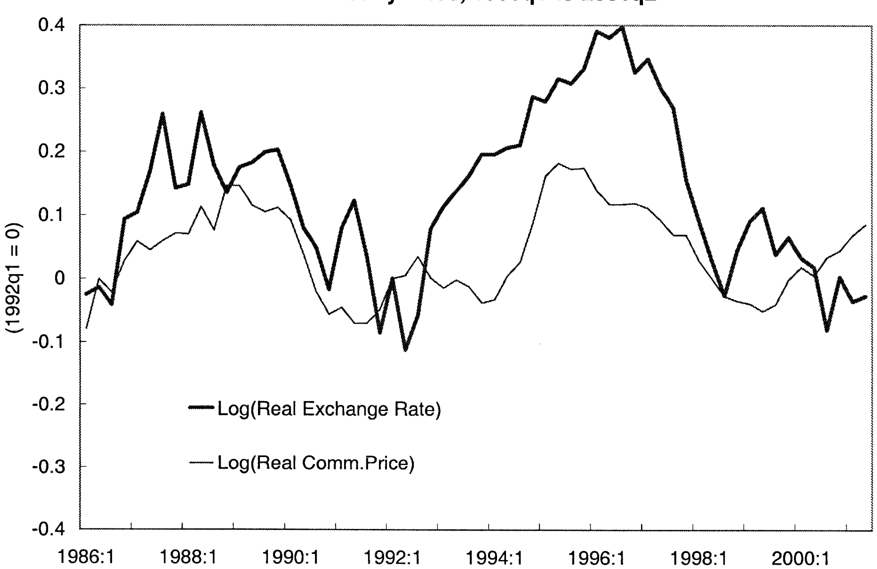

Fig. 7a: Australian-US Real Exchange Rate and Traded vs. Non-Traded Relative Productivity Differential

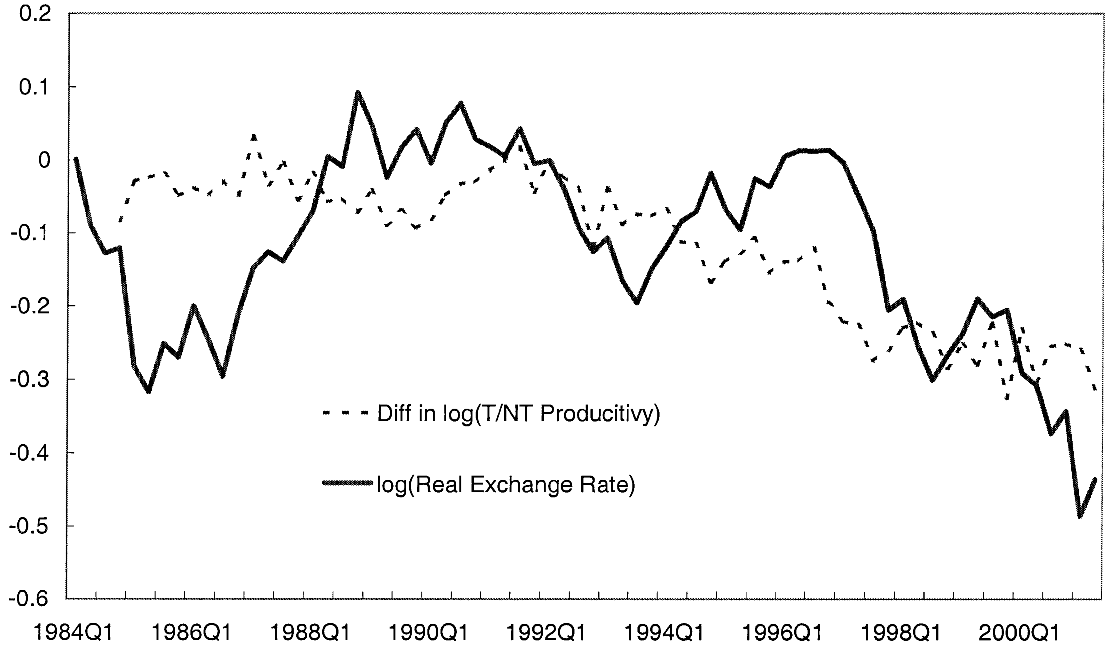

Fig. 7b: NZL-US Real Exchange Rate and Traded vs. Non-Traded Relative Productivity Differential