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THE COMMODITY-CURRENCY VIEW OF THE AUSTRALIAN DOLLAR: A MULTIVARIATE COINTEGRATION APPROACH

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Using Australian quarterly data from the post-float period 1984:1-2003:1 and a partial system, we identify and estimate two cointegrating relations, one for the interest-rate differential and the other for the nominal exchange rate. Our estimate of the long-run elasticity of the exchange rate with respect to commodity prices is 0.939, which strongly supports the widely held view that the floating Australian dollar is a 'commodity currency'. We also find that the PPP and UIP cannot be rejected so long as commodity prices are included in the cointegrating relations. Our model outperforms the random walk model in forecasting the exchange rate in the medium run.

JEL classification codes: F31, F41
Key words: Australian dollar, commodity currency, cointegration

I. Introduction

In the Australian setting, extraneously determined terms of trade have long been recognized as a variable playing a central role in influencing the country's economic outcomes (Salter, 1959; Swan, 1960; Gregory, 1976).

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Since the floating of the Australian dollar ($A) in December 1983, attention has focused more sharply on how terms of trade volatility projects into volatility of the exchange rate (nominal and real) and how this impinges on Australian competitiveness, macroeconomic stability, and resource allocation.

In their search for an empirical counterpart of such a link, Blundell-Wignall et al. (1993) postulated that a cointegrating relationship exists between the real exchange rate, the terms of trade, the long-term real interest differential, and the ratio of net foreign assets to GDP. Their key finding is that a 10% improvement in the Australian terms of trade is associated with a real appreciation of the $A by about 8%. Subsequent studies also employ cointegration analysis to estimate the terms-of-trade elasticity of the real or the nominal exchange rate (Gruen and Wilkinson, 1994; Koya and Orden, 1994; Fisher, 1996; Karfakis and Phipps, 1999). The estimates of these two elasticities often exceed unity. Thus, using the $A/$US exchange rate and US-dollar based terms of trade data, Fisher (1996, Table 2) estimates the two elasticities as 1.29 and 1.45, respectively, a result which is similar to that reported by Koya and Orden (1994, Table 3).

A common point of departure of these studies is that the cointegrating relationship involves terms of trade. In Australia, however, the terms of trade themselves correlate highly with phases of the world commodity price cycle, confirming the overpowering influence of fluctuating commodity prices as the mechanism that delivers external shocks to the exchange rate. Thus, our point of departure is to test if a direct link exists between a commodity-price index and the nominal value of the $A. Such a link lies at the heart of the well-known “commodity currency” view of the $A (Clements and Freebairn, 1991; Hughes, 1994; Chen, 2002; Chen and Rogoff, 2003). According to this view, the $A appreciates (depreciates) in both nominal and real terms when the prices of certain commodities exported by Australia, e.g., coal, metals, and other primary industrial materials, rise (fall) in international markets.

If it is indeed true that economic agents react to commodity prices in their purchases or sales of the $A in the foreign exchange market, it is difficult to sustain the notion that they would be doing so on the basis of terms of trade data, which are available quarterly and then only with a long publication lag. By contrast, price quotations for the principally traded commodities are published daily in the financial press. Likewise, the nominal trade-weighted exchange rate index (TWI) is made available daily by the Reserve Bank of Australia (RBA), while its inter-day movements may be inferred from continuing market developments. It is this wealth of information that underlies the data set that forms the basis for this study.

We begin our search for a long-run equilibrium nominal exchange-rate equation by adopting a four-dimensional VAR model (section II). After discussing the data (section III), we use a standard cointegration analysis (section IV) and find two cointegrating relations, so we are confronted with an identification problem, which is generally difficult to deal with. The papers cited earlier do not address this problem, because they either find empirically or simply assume only one cointegrating relation. Here, we address the identification problem and test a number of hypotheses in the context of a partial system. Our estimate of the long-run elasticity of the exchange rate with respect to commodity prices is 0.939 and statistically not different from unity, which strongly supports the commodity-currency hypothesis. We also construct a parsimonious forecasting model (section V). Section VI concludes the paper.

II. The Economic and the Statistical Model

Let e_{12} be the logarithm of the \$A price of one unit of foreign currency; $p_{12} = \ln(p_1) - \ln(p_2)$; and $i_{12} = \ln(1+i_1) - \ln(1+i_2) \approx i_1 - i_2$ (for small values of i_1 and i_2), where p_1, p_2, i_1, and i_2 are Australian and foreign price levels and interest rates, respectively. We assume that uncovered interest parity (UIP) holds approximately, i.e.,

$$i_{12,t} \approx E_t(e_{12,t+1}) - e_{12,t}, \tag{1}$$

where $E_t(.)$ is a conditional expectation formed at time t. We also assume that the expected long-run value of the exchange rate is determined according to the equation,

$$E_t(e_{12,t+1}) = \omega_1 p_{12,t} + \omega_2 cp_t, \tag{2}$$

where cp_t is a commodity-price index and ω_1 and ω_2 are coefficients used by the forecaster. Equation (2) is a modification of Equation (1) of Karfakis and

Phipps (1999), which uses p_{12} and terms of trade as determinants of $E_t(e_{12,t+1})$. As we discussed in the Introduction, this modification seems preferable when Australian data are used. Substituting (2) into (1) and assuming that at least two of the four variables e_{12}, i_{12}, p_{12}, and cp are integrated of order one, I(1), whereas the others are stationary, I(0), it becomes evident that the model can have empirical content only if a linear combination of these variables is stationary:

$$e_{12,t} + \beta_1 i_{12,t} + \beta_2 p_{12,t} + \beta_3 cp_t \sim \mathrm{I}(0). \tag{3}$$

Note that the restriction imposed by purchasing power parity (PPP) is $\beta_2 = -\omega_1 = -1$ and that imposed by UIP is $\beta_1 = 1$ (see Juselius, 1995, p. 214).

Based on condition (3), we use in our cointegration analysis the following set of stochastic variables: $\mathbf{z}_t' = \{e_{12}, i_{12}, p_{12}, cp\}_t$. This definition of $\mathbf{z}_t$ combined with our set of dummy and dummy-type variables $\mathbf{D}_t$ (see section III) minimizes the problems of model misspecification and identification of the long-run coefficients.[1] We begin by considering a four-dimensional vector autoregressive (VAR) model, which can be written in the form of a vector error-correction model (VECM) as follows:

$$\Delta\mathbf{z}_t = \mathbf{\Gamma}_1\Delta\mathbf{z}_{t-1} + \mathbf{\Gamma}_2\Delta\mathbf{z}_{t-2} + \ldots + \mathbf{\Gamma}_{k-1}\Delta\mathbf{z}_{t-(k-1)} + \mathbf{\Pi}\tilde{\mathbf{z}}_{t-1} + \mathbf{\Psi}\mathbf{D}_t + \boldsymbol{\varepsilon}_t, \tag{4}$$

where $\tilde{\mathbf{z}}_t' = \{e_{12}, i_{12}, p_{12}, cp, t\}_t$ and t = time trend, so we allow for trend stationarity in the cointegration relations.[2] We choose the lag length (k) and

[1] We also experimented with alternative models by augmenting the forecasting Equation (2) and the vector $\mathbf{z}_t$ to include the logarithm of one or more of the following variables: real GDP (or unemployment), current-account deficit, and net foreign assets as a share of GDP in an attempt to capture the possible influence of Australia's rising foreign debt. In most cases, the results that are of interest did not change substantially, but the diagnostic tests consistently worsened and the economic identification (in the sense of Johansen and Juselius, 1994) of the cointegrating vectors and of the adjustment coefficients became difficult. Also, in most cases the number of cointegrating vectors appeared to increase by at least one, thus requiring additional identifying restrictions. Finding such restrictions had to be more or less arbitrary, however.

[2] The time trend in the cointegrating relations accounts for the level of our ignorance regarding variables that influence $\mathbf{z}_t$ systematically, but are not present in our cointegrating relations.

the variables in the vector $\mathbf{D}_t$ so as to make the errors $\boldsymbol{\varepsilon}_t$ Gaussian white noise. The matrix $\mathbf{\Pi}$ is 4×5 and has rank r, where $0 \leq r \leq 4$. As is well known, cointegration arises when $1 \leq r \leq 3$, in which case we write $\mathbf{\Pi} = \boldsymbol{\alpha\beta}'$, where $\boldsymbol{\beta}$ is a 5×r matrix whose columns are the r cointegrating vectors and $\boldsymbol{\alpha}$ is a 4×r matrix containing the "speed-of-adjustment coefficients."

III. The Data

We use quarterly data from the post-float period 1984:1-2003:1. The end period for estimation is 2001:4, since we keep the last five observations to assess the model's forecasting performance. Here, e_{12} = logarithm of the trade-weighted nominal exchange rate (Reserve Bank of Australia, RBA); i_1 = Australian 90-day bill rate (bank accepted bills, quarterly averages of monthly figures, RBA); i_2 = 90-day Eurodollar rate (quarterly averages of monthly figures, RBA); p_1 = the Australian Consumer Price Index (CPI), 1995 = 100 (Australian Bureau of Statistics); p_2 = OECD-Total CPI, 1995 = 100 (OECD); and cp = logarithm of the index of commodity prices (all groups) measured at external prices (SDRs), 1994-95 = 100 (RBA).[3]

In our effort to satisfy the assumption of independently and normally distributed errors in Equation (4), preliminary analysis suggested that we define the vector $\mathbf{D}_t$ as,

$$\mathbf{D}_t' = \{1, \Delta po_t, \Delta po_{t-1}, \Delta po_{t-3}, d84_t\}. \tag{5}$$

The first element of $\mathbf{D}_t$ is unity, which means that we include a constant term in each equation of the system (4), implying that we allow for trend stationarity in the variables. The next three elements of $\mathbf{D}_t$ are current and lagged percentage changes in the price of crude oil (UN, Monthly Bulletin of Statistics).[4] Finally, we define the dummy variable *d84* as *d84* = 1 for

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[3] The RBA calculates the commodity price index in three different ways, using the US dollar, the Australian dollar, and the SDR unit as the currency denomination. Because not all Australian commodity exports are traded for US currency, we employ the SDR-based series as the broadest price measure available.

[4] Following Hansen and Juselius (1995, pp. 17-18), we treat Δpo_t as a dummy-type variable, since it is assumed to be both weakly exogenous for $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ and absent from the

1984:1-1985:3 (*d84* = 0 otherwise) to reflect the switch from the previous highly regimented "administered system" (a form of crawling peg) to the new floating rate regime. The float was accompanied by the abolition of the still extant system of war-time exchange controls, allowing economic agents virtually complete freedom in managing their foreign transactions. Because the market took time to grasp the new *modus operandi* of the system, *d84* may be termed a learning-process dummy.

IV. Cointegration and Error-correction Modeling

We begin our econometric analysis by testing for unit roots and by determining the lag length of the VAR in levels.[5] As is well known, unit-root tests have low power, so we use several of them, namely the augmented Dickey-Fuller and Phillips-Perron tests and those suggested by Perron (1989), Kwiatkowski et al. (1992), and Hylleberg et al. (1990). We conclude that each of the four variables in z_t has a non-seasonal unit root, but is seasonally stationary. Next, using Sims's likelihood-ratio (LR) test and Johansen's (1995, p. 21) advice that "it is important to avoid too many lags," we choose a lag length of k = 3.[6]

Then, we estimate Equation (4) by the Johansen procedure, as implemented by the computer program CATS in RATS (Hansen and Juselius, 1995). The multivariate tests computed by this program for the hypothesis of white noise errors against autocorrelation of order 1, 4, and 17 do not reject this crucial hypothesis upon which the following methods are based (see Johansen, 1995, p. 21), since their p-values are 0.53, 0.20, and 0.35, respectively. The tests

cointegration space. If the variables Δpo_{t-i}, i = 0, 1, 3, are not included in $\mathbf{D}_t$, then the residuals are weakly correlated with them. (The highest of these correlations is 0.2 in absolute value with a t-ratio of 1.67. When these variables are included in $\mathbf{D}_t$, however, these correlations reduce to zero, as expected.)

[5] Here and in what follows we report only the conclusions from our tests. Detailed results are available upon request from the first author.

[6] At the 1% level, the LR test suggests that we choose k = 6. Such a choice makes economic identification more difficult, however. Thus, since the test suggests that the fourth and the fifth lag are not significant even at the 5% level, we follow Johansen's advice and choose k = 3.

reject the hypotheses of normality and of no ARCH, however, even at the 1% level. In particular, there is a strong ARCH effect in the equation for Δi_{12}, which occurs because the residuals of this equation for the quarters 1985:2-1985:4 and 2001:3-2001:4 are relatively large. Following Harris (1995, p. 86), we solve this problem by introducing an additional dummy, *darch*, which takes on the value of one for the above quarters and zero otherwise. As for the normality assumption, it will be satisfied at the 5% level when we consider a partial system (see below), so its failure at this stage should not cause a great concern, especially because it is not crucial for the Johansen procedure (see Gonzalo, 1994).

Thus, we proceed to the next step, which is the determination of the cointegration rank, r. This step is crucial, since our results will be conditional on the choice of r (Hansen and Juselius, 1995, p. 8). Following Johansen and Juselius (1990, 1992), Hansen and Juselius (1995), and Harris (1995, pp. 86-92), we use well-known statistical, graphical, and theoretical criteria, all of which suggest that r = 2.[7] Thus, we set r = 2. We normalize the first cointegrating relation by i_{12} and the second by e_{12}, because it looks like an exchange rate equation.

We then perform the following two tests. First, a LR test fails to reject the PPP restriction in both cointegrating relations (p-value = 0.23). Second, we test for weak exogeneity of the variables using a two-tailed t-test for each of the hypotheses H_0: $\alpha_{ij} = 0$, i = 1, ..., 4, j = 1, 2. In each of the equations for Δe_{12} and Δi_{12} at least one of these two hypotheses is rejected even at the 1% level, whereas in the equations for Δp_{12} and Δcp none of the two hypotheses is rejected even at the 10% level (t-ratios: 0.16 and 0.03 in the equation for Δp_{12} and 0.65 and 1.50 in the equation for Δcp). Thus, out-of-equilibrium values of i_{12} and e_{12} do not cause changes in p_{12} and in cp, so the latter can be taken as weakly exogenous variables for β and the remaining α's. A LR test of the joint hypothesis H_0: $\alpha_{31} = \alpha_{32} = \alpha_{41} = \alpha_{42} = 0$ supports this result (p-value = 0.67). Weakly exogenous behavior for Δcp might have been expected, since, on the whole, Australia appears to be a price taker in world markets for most

[7] Our theoretical criterion is to count the number of eigenvectors that satisfy the PPP restriction.

of its exported commodities.[8] Such behavior for Δp_{12} is somewhat surprising, however, and may be attributed to short-run price rigidity or to the use of formal inflation targeting since 1993 (see Zettelmeyer, 2000, p. 10). In what follows, we treat Δp_{12} and Δcp as weakly exogenous variables and proceed with the partial system

$$\Delta \mathbf{y}_t = \breve{\mathbf{\Gamma}}_0 \Delta \mathbf{x}_t + \breve{\mathbf{\Gamma}}_1 \Delta \mathbf{z}_{t-1} + \boldsymbol{\alpha}\boldsymbol{\beta}'\breve{\mathbf{z}}_{t-1} + \breve{\mathbf{\Psi}}\breve{\mathbf{D}}_t + \breve{\boldsymbol{\varepsilon}}_t, \tag{6}$$

where $\mathbf{y}_t' = \{e_{12}, i_{12}\}_t$, $\mathbf{x}_t' = \{p_{12}, cp\}_t$, and $\breve{\mathbf{D}}_t$ is the vector $\mathbf{D}_t$ augmented to include the dummy variable *darch* (defined earlier). Again, we find r = 2 cointegrating vectors.

Now the p-values of the multivariate tests for white noise errors against autocorrelation of order 1, 4, and 17 are 0.53, 0.10, and 0.60, respectively, whereas that for normality is 0.06. The univariate tests indicate that the normality assumption cannot be rejected in the equation for Δe_{12} (p-value = 0.43), but can be rejected in the equation for Δi_{12} (p-value = 0.004). Finally, the hypothesis of no ARCH (against ARCH of order 3) cannot be rejected now for either equation (p-values: 0.72 and 0.44). Thus, since normality is not crucial for the Johansen procedure (Gonzalo, 1994), we consider Equation (6) adequate and use it to test some hypotheses, which are either theoretically interesting or useful as identifying restrictions.

First, in accordance with previous studies, we find that ignoring commodity prices results in a rejection of PPP, even when it is combined with UIP (p-value = 0.00). Taking these prices into account, however, we cannot reject the hypothesis that PPP holds in both cointegrating relations (p-value = 0.35); neither can we reject the joint hypothesis that PPP holds in both cointegrating relations and UIP holds in the second (p-value = 0.15). Second, using a LR test, we find that the variables i_{12}, cp, and t form a cointegrating relation ($\chi_1^2 = 0.07$, p-value = 0.79). The results of the last two tests are used below as identifying restrictions.

Johansen and Juselius (1994, p. 15) provide a necessary and sufficient condition for generic identification, which is given by the following inequality:

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[8] See Chen and Rogoff (2003, pp. 136 and 147). An example where Australia has market power is the case of wool. See Clements and Freebairn (1991, p. 4).

$$r_{i.j} = rank(\mathbf{R}_i\,'\mathbf{H}_j) \geq 1, \quad i \neq j. \tag{7}$$

Here, $\mathbf{R}_i$ and $\mathbf{H}_i$ are, respectively, 5×k_i and 5×(5-k_i) design matrices of full rank and with known coefficients, such that $\mathbf{R}_i\,'\mathbf{H}_i = \mathbf{0}$, $\mathbf{R}_i\,'\boldsymbol{\beta}_i = \mathbf{0}$, or equivalently $\boldsymbol{\beta}_i = \mathbf{H}_i\boldsymbol{\varphi}_i$ for some (5-k_i)-vector $\boldsymbol{\varphi}_i$, where k_i is the number of restrictions imposed on the i-th cointegrating relation. Because of the finding that the variables i_{12}, *cp*, and *t* form a cointegrating relation and because no violence was done to the data when UIP was imposed on the second cointegrating vector, we specify the design matrices as follows:

$$\mathbf{H}_1 = \begin{bmatrix} 0 & 0 & 0 \\ 1 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{bmatrix}, \quad \mathbf{R}_1 = \begin{bmatrix} 1 & 0 \\ 0 & 0 \\ 0 & 1 \\ 0 & 0 \\ 0 & 0 \end{bmatrix}, \quad \mathbf{H}_2 = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}, \quad \mathbf{R}_2 = \begin{bmatrix} 1 \\ -1 \\ 0 \\ 0 \\ 0 \end{bmatrix}. \tag{8}$$

Thus, we impose $k_1 = 2$ restrictions on the first cointegrating relation (that the coefficients of e_{12} and p_{12} are both zero) and $k_2 = 1$ restriction on the second relation (that UIP holds). These restrictions, which will be tested below, satisfy condition (7), since $r_{1.2} = 2$ and $r_{2,1} = 1$. Thus, imposing them results in an identifiable and, therefore, estimable model.

Table 1 reports coefficient estimates of the two cointegrating relations, their standard errors, and univariate diagnostic tests. Note that the p-values of the multivariate tests for white noise errors against autocorrelation of order 1, 4, and 17 are 0.52, 0.10, and 0.60, respectively, whereas that for normality is 0.06. Also note that a LR test of the restrictions imposed by $\mathbf{H}_1$ *and* $\mathbf{H}_2$ gives $\chi^2_1 = 0.07$ with a p-value of 0.79. Since this result is identical to that obtained when we tested only the restrictions imposed by $\mathbf{H}_1$ (that the variables i_{12}, *cp*, and *t* form a cointegrating relation), it follows that the restriction imposed by $\mathbf{H}_2$ is just identifying, and only $\mathbf{H}_1$ imposes restrictions on the cointegration space (see Hansen and Juselius, 1995, p. 43). Further adequacy and structural stability tests reveal no strong evidence against the model, so the estimates of Table 1 are deemed usable. We discuss the most important of them.

Begin with the second cointegrating relation, our exchange-rate equation. The coefficient of *cp*, which is the elasticity of e_{12} with respect to *cp*, is 0.939

Table 1. Estimates of α and β and Univariate Diagnostic Tests from an Identified VECM

	Cointegrating vectors, $\hat{\beta}$					Adj. coef., $\hat{\alpha}$, and univariate diag. tests	
	e_{12}	i_{12}	p_{12}	cp	t	Equation for Δe_{12}	Equation for Δi_{12}
$\hat{\beta}_1$	0	1.000	0	-0.118	0.001	-0.259	-0.651
(s.e.)	---	---	---	(0.018)	(0.000)	(0.273)	(0.063)
$\hat{\beta}_2$	1.000	1.000	-1.567	0.939	-0.004	-0.442	-0.085
(s.e.)	---	---	(0.319)	(0.124)	(0.001)	(0.089)	(0.021)
ARCH(3)	---	---	---	---	---	1.30	2.67
JB	---	---	---	---	---	1.69	11.16

Note: The statistics ARCH(3) and JB (Jarque-Bera test for normality) are approximately distributed as χ^2_3 and χ^2_2, respectively.

with a standard error of 0.124, so this elasticity is statistically not different from unity. Thus, in the long-run, a 10% increase in the commodity-price index causes an appreciation of the \$A by almost 10%.[9] This estimate provides strong support for the well known "commodity currency" story, here modified for the role of the short-term interest differential. Freebairn (1991, pp. 23-28) explains its profound implications for the Australian economy, which can be understood by considering a world commodity-price boom, a positive demand shock that causes a nominal and a real appreciation of the \$A. In addition to its macroeconomic effects, this shock will influence the profitability of the various sectors in Australia differently, thus causing a reallocation of resources. Freebairn's Figures 6-8 suggest that the closer this elasticity is to unity the weaker will be the (positive) effects on the export and the non-traded sectors, since the appreciation of the \$A will partially offset the first-round effects of

[9] Note that the sign of this elasticity is in fact negative, since *cp* should be thought of as a right-hand-side variable in the exchange rate equation. In Table 1, this elasticity has a positive sign, because it is reported as an estimate of the coefficient β_3 in Equation (3).

the commodity-price boom, and the stronger will be the (negative) effect on the import-competing sector.

The estimate 0.939 of this elasticity is similar to that reported by Chen (2002, Table 1A) for the case of the \$A/\$US exchange rate, 0.92, but is higher than Chen's estimates for other currency pairs, and is also higher than the "conventional wisdom" estimate, 0.5 (see Clements and Freebairn 1991, p. 1). Note that, to our knowledge, Chen (2002) and Chen and Rogoff (2003) are the only other papers that examine in depth statistically the effect of commodity prices on the exchange rate, with the former focusing on the nominal and the latter on the real rate. Both of these papers assume, however, that there exists only one cointegrating vector and estimate it by dynamic OLS (DOLS), which, under this assumption, is asymptotically equivalent to Johansen's method. But if there exist two cointegrating vectors and the one estimated has larger variance than the one ignored, then Monte Carlo evidence suggests that single-equation methods (such as DOLS) are inappropriate (see Maddala and Kim, 1998, pp. 183-184). Both of these conditions hold in our data,[10] so Johansen's method is preferable.

Having identified and estimated a long-run equilibrium exchange-rate equation, we are able now to test the absolute PPP restriction in Equation (3), $\beta_2 = -1$. Using a Wald test, we cannot reject absolute PPP at the 5% level, since $\{[-1.567-(-1)]/0.319\}^2 = 3.16 < \chi^2_{1,0.05} = 3.84$.

Next, consider the first cointegrating relation, our long-run equilibrium equation for the interest-rate differential. The coefficient of *cp*, 0.118 (t-ratio = 6.56), means that in the steady-state a 10% increase in the commodity-price index will push up domestic (relative to foreign) short-term interest rates by 0.0118 of a percentage point.[11] For a commodity-price boom stimulates domestic income and expenditure and raises expectations of inflation.

We now turn to the speed-of-adjustment coefficients. The first error-

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[10] That is, (1) we have two cointegration vectors, and (2) the residuals from the exchange rate equation have larger variance than those from the equation for the interest-rate differential.

[11] This interpretation is based on the following approximation: if $y = \beta \ln(x)$, then $\Delta y \approx \beta(\Delta x)/x$.

correcting term (*ecm1*) is interpreted as a disequilibrium interest-rate differential. It is statistically significant only in the equation for Δi_{12}, with a coefficient of -0.651 (t-ratio = -10.41), which implies that about 65% of a disequilibrium interest-rate differential is removed in one quarter.

The second error-correcting term (*ecm2*) is interpreted as a disequilibrium exchange-rate. Its coefficient in the equation for Δe_{12} is –0.442 (t-ratio = -4.95), which implies that about 44% of a disequilibrium value of the $A is removed in one quarter. In the equation for Δi_{12}, the coefficient of *ecm2* is -0.085 (t-ratio = -4.13). A possible interpretation of the sign of this coefficient is as follows. Assume an excessive depreciation of the $A in the previous quarter (i.e., a large value of *ecm2* in quarter t-1). If this should raise market expectations that a reversal in the exchange rate is imminent, speculative capital inflows would provide buying support for the $A in quarter t while also exerting a downward pressure on the domestic interest rate and bringing about a correction in the interest diferential.[12]

Finally, note that in the context of the partial VECM, relative PPP cannot be rejected, since the coefficient of Δp_{12} in the equation for Δe_{12} is 0.996 with a standard error of 0.56. Thus, in the next section, where we estimate a parsimonious model, we impose the restriction that this coefficient is unity. Since in the next section we report and discuss the estimates of the short-run parameters of interest obtained from the parsimonious model, here we do not do so for the remaining estimates of the partial VECM.

V. A Parsimonious Model

A number of short-run coefficients in the estimated partial VECM are statistically insignificant. Thus, we now construct a parsimonious VECM

[12] Other scenarios are possible. If the RBA should believe that the out-of-equilibrium value of the exchange rate is not self-correcting, or that a further depreciation is in prospect, it might try to aid the $A by raising the domestic interest rate. If the interest adjustment in quarter t-1 should in turn prove excessive, it would call for a correction in quarter t. Note, however, that such a policy action is treated here as an exogenous shock captured by the error term of the equation for Δi_{12} in quarter t-1 (see Zettelmeyer, 2000, p. 23).

that incorporates relative PPP,[13] in that the dependent variable of the first equation is $RPPP = \Delta e_{12} - \Delta p_{12}$. We estimate each equation of this VECM separately by a least-squares method that is robust to autocorrelation. Table 2 reports the results. There is no autocorrelation at the 5% level. We use this model to forecast the exchange rate and to estimate some short- or medium-run dynamic effects of changes in the weakly exogenous variables, Δcp and Δp_{12}, on the endogenous variables, Δe_{12} and Δi_{12}.

First, consider a *ceteris paribus* increase in the rate of increase of the commodity-price index by 1 percentage point. This shock will cause the rate of appreciation of the \$A to rise by 0.67 of 1 percentage point during the same quarter; by 0.44 of 1 percentage point in two quarters time, since -0.67 + 0.23 = -0.44; and by 0.37 of 1 percentage point in four quarters, since (-0.67 + 0.23)/(1 + 0.19) = -0.37. It will also cause the rate of increase of the interest-rate differential to rise by about 0.06 of 1 percentage point during the same quarter.

Second, consider a *ceteris paribus* increase in the rate of increase of the inflation differential, Δp_{12}, by 1 percentage point. This will cause the rate of depreciation of the \$A to rise by 1 percentage point during the same quarter, since the model incorporates relative PPP; and if it persists for another quarter, it will eventually cause Δi_{12} to rise by a total of 1.15 (= 0.56 + 0.59) percentage points in two quarters time.

Note also that the three estimates of the speed-of-adjustment coefficients produced by this model are all somewhat smaller (in absolute value) than those produced by the Johansen procedure. Although the difference is not statistically significant in the case of the equation for Δi_{12}, in which these coefficients are -0.59 (versus -0.651) and -0.083 (versus -0.085), the question still arises that, by dropping the insignificant terms from the partial VECM, in our effort to construct a parsimonious model, we may have introduced a bias.

According to the recent literature, an "acid test" that exchange rate models

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[13] Recall from the end of the previous section that relative PPP could not be rejected. Also, note that we do not construct a simultaneous equations model, because the contemporaneous correlation between the endogenous variables Δe_{12} and Δi_{12} is only 0.28 and because in doing so a number of new issues arise, namely, re-specification, identification, and choice of instruments.

Table 2. Estimates from a Parsimonious VECM (t-ratios in parentheses)

Dep. Var.	Const. term	Δcp_t	Δcp_{t-2}	Δi_{12t-1}	Δp_{12t-1}	Δp_{12t-2}	$RPPP_{t-4}$	Δpo_t	Δpo_{t-1}	Δpo_{t-3}	$d84_t$	$darch$	$ecm1_{t-1}$	$ecm2_{t-1}$
$RPPP_t$	2.61	-0.67	0.23	—	—	—	-0.19	-0.0005	—	—	—	—	—	-0.28
	(4.5)	(-6.3)	(2.1)				(-2.5)	(-1.9)						(-4.5)
Δi_{12t}	0.50	0.06	—	0.19	0.56	0.59	—	—	-0.0002	-0.0002	-0.03	0.02	-0.59	-0.08
	(2.2)	(3.0)		(2.1)	(3.7)	(3.3)			(-2.0)	(-1.9)	(-3.2)	(6.4)	(-5.1)	(-3.0)

Note: The values of R^2 are 0.48 and 0.60, respectively. The ranges of the p-values for modified LM tests for first- to fourth-order autocorrelation as well as seasonal autocorrelation are 0.330 - 0.935 and 0.052 - 0.923, respectively (see Godfrey, 1988, pp. 178-179).

must pass in order to be deemed worthy of consideration is forecasting out-of-sample better than a random walk. Using our parsimonious model, we generate one-step ahead dynamic forecasts for e_{12} for five out-of-sample quarters, namely 2002:1-2003:1, and calculate Theil's U-statistic (see MacDonald and Nagayasu, 1998, p. 98). For the one-, two-, . . ., five-quarter time horizons, we find the following values of U: 1.03, 0.38, 0.57, 0.59, and 0.48. With the exception of the one-quarter horizon, all of the other values of U are well below unity, so the model outperforms the random walk model in forecasting the values of e_{12} in the medium run.

VI. Summary and Conclusions

Using Australian quarterly data from the post-float period 1984:1-2003:1, we find two steady-state relationships, one for the interest differential and the other for the nominal exchange rate, and hence two error adjustment mechanisms, suggesting that the transition from one equilibrium to another is attended by interactions between goods markets and assets markets. An external disturbance, such as a commodity price shock, will set off an adjustment mechanism causing both the exchange rate and the interest differential to adjust to their new equilibrium levels – an interpretation which accords well with the rapid integration of Australian markets and overseas markets which followed the float. Our estimate of the steady-state elasticity of the nominal exchange rate with respect to the commodity price index is 0.939, whereas that of the short-term dynamic effect is 0.67.

According to these estimates, a *ceteris paribus* increase in commodity prices, which improves the Australian terms of trade, boosts export income, and generates a trade surplus, will stimulate foreign demand for Australian dollars and will initially cause the $A to appreciate almost proportionately. This nominal appreciation may produce a deflationary effect in the real sector, which, unless offset by policy, is apt to alter the economic agents' perceptions of what the financial variables in the system (interest differential and trade-weighted exchange rate) "ought to be" in the changed circumstances. If the market sentiment should be that the initial currency windfall has overvalued the $A, the reaction would be to sell $A, triggering the error adjustment mechanism, which eventually propels the actual rate towards its steady-state.

Should the initial nominal appreciation for some reason "undervalue" the \$A in terms of the commodity fundamentals, the convergence would be in the opposite direction. We estimate that about 44% of the divergence between the actual and the steady-state value of the exchange rate will be eliminated with a lag of one quarter.

Although we fail to reject PPP and UIP, so long as commodity prices are included in the cointegrating relations, note that the PPP relation is inherently difficult to capture in a study of this type, for domestic price developments will not be uninfluenced by substantial shifts in domestic monetary and fiscal policies, and these are not explicitly accounted for in our model. Thus, it would seem hazardous to attempt to search for a cointegrating relationship for the *real* exchange rate, unless we can adequately account for the possible influences of policy changes on the price level. This caveat is important, for it was the highly restrictive official response to adverse external developments in 1988-1989, which appears to have ushered the new low inflationary environment in Australia that persisted throughout the 1990s.

As a final check, although in the case of the one-quarter horizon our model does not outperform the naive random walk in forecasting the exchange rate, it does so in the two-, three-, four-, and five-quarter time horizons. Thus, all things considered, our model does not seem to be an unreasonable approximation of the true mechanism underlying the observed behavior of the floating Australian exchange rate.

The most important implication of our findings can be stated as follows. Since about 80% of the Australian merchandise exports consist of commodities at various stages of processing, and since the exchange rate in its steady state moves almost one-for-one with world commodity prices, the cyclical path in these prices maps closely into cyclical behavior of the nominal effective exchange rate, with powerful implications for the international competitiveness of Australia's elaborately transformed exports and import competing goods. While it is true that this mechanism has helped produce an economic environment in Australia which is far less prone to the inflationary excesses experienced under the previous regime of administered exchange rates, it cannot be said that it 'protects' Australians from the instability which is inherent in the workings of the international commodity markets. It only transfers that instability into another domain.

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