The monetary approach to exchange rates and the behaviour of the Canadian dollar over the long run

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Using Canadian–US dollar data this paper examines the question of whether recent positive findings with regard to purchasing power parity carry over to the monetary approach to exchange rates. The evidence provides strong support for the long-run monetary model of exchange rates. At the same time, it provides indirect evidence in favour of long-run purchasing power parity between the US dollar and the Canadian dollar during the sample period.

I. INTRODUCTION

A decade ago, the consensus view in international finance was that purchasing power parity (PPP) was of little use empirically and that models of exchange rate behaviour that relied upon it were of little or no practical relevance.[1] In the past five years, however, such assessments have had to be tempered as studies have increasingly shown that as a long-run equilibrium condition, PPP has, in fact, been a useful first approximation, both historically and, according to a number of recent studies, under the current float.[2]

The purpose of this paper is to examine whether these more favourable findings with regard to purchasing power parity carry over to the monetary approach to exchange rates. Evidence to date on this subject has been rather scanty and somewhat mixed. The results reported here for the Canadian–US dollar rate, do however, support the monetary approach to exchange rates, particularly as a long-run equilibrium relation.[3]

What makes these results of particular interest is the close links between monetary policies in the two countries. One of the criticisms levied against the studies of purchasing power parity that have proliferated over the past decade is that most are for countries in which policies at one time or another have been widely different. Real effects on exchange rates though perhaps substantial in absolute terms will necessarily prove small in relative terms. Such studies, it is claimed, will as a result overstate the degree to which real exchange rates have been stable, and in particular reject the unit-root hypothesis when real exchange rates do in fact have economically meaningful permanent components.

Empirical results are reported in the third section of the paper. In the second, the basic monetary approach model is derived and the existing empirical work reviewed briefly both on the monetary approach to exchange rates, and, because it is closely related, on purchasing power parity too. The last section of the paper contains conclusions along with several suggestions with regard to future work in this area.

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[1] See, for example, the survey articles written near the end of the 1980s by Dornbusch (1987), Frankel and Meese (1987), and Meese, (1990).

[2] See the references cited in notes 4 and 5 below.

[3] Choudhry and Lawler (1997) who examined the validity of the monetary model of exchange rate determination as an explanation of the Canadian dollar–US dollar relationship over the Canadian float 1950–1962 report similar conclusions.

II. THEORY AND PREVIOUS EVIDENCE

Theory

Underlying the monetary approach to exchange rates (MAER) are two basic building blocks, the quantity theory of money and purchasing power parity (PPP). To derive the model in its simplest form, we start with the following three equations:

$$s_t = p_t - p_t^* \tag{1}$$

$$p_t = m_t - \eta(y_t) = \lambda i_t \tag{2}$$

$$p_t^* = m_t^* - \eta^*(y_t^*) + \lambda^* i_t^* \tag{3}$$

where s is the nominal exchange rate (defined as domestic price of a unit of foreign currency), p and p^* are the domestic and foreign price levels, m and m^* are the domestic and foreign money supplies, y and y^* are domestic and foreign real income, i and i^* are domestic and foreign nominal interest rates, 0, 0^*, λ and λ^* are parameters of the two countries' money demand functions, and where all variables other than the interest rates are in logarithmic form. Combining Equations 1–3 gives:

$$s_t = [m_t - \eta(y_t) + \lambda i_t] - [m_t^* - \eta^*(y_t^*) + \lambda^* i_t^*] \tag{4}$$

Here the nominal exchange rate is seen as determined solely by the contemporaneous excess supplies of money in the two countries. Underlying Equation 4 is the intuitively appealing idea that countries that follow relatively expansive monetary policies see their exchange rates depreciate, while countries that follow relatively restrictive policies experience the opposite. It is, therefore, best viewed as a long-term equilibrium relationship although in early empirical applications it was at times applied to differenced data over relatively short time periods. In the tests of the model, this equilibrium property is exploited.

Because all of the variables in Equation 4 generally have been found to be nonstationary in levels, or $I(1)$, but stationary in first differences, the equilibrium posited by Equation 4 can only occur if s_t and the right-hand-side variables in Equation 4, the 'fundamentals' influencing nominal exchange rates, form a cointegrating relationship. Only then will their linear combination be stationary, or $I(0)$, and short-run deviations of s_t from the value consistent with

the values of the fundamentals become zero in the long run. To test the long-run implications of the model cointegration between s_t and the fundamentals are tested for.

Previous results

The early empirical implementation of the MAER met with a good deal of success (e.g. Frenkel, 1976; Bilson, 1978). By the mid-1980s, however, that changed. Even versions of the MAER that were much less restrictive than Equation 4 performed quite poorly. The estimated parameters of these equations usually were inconsistent with theory and at times even of the wrong sign. The models, moreover, explained little of the actual variation in nominal exchange rates (Frankel, 1984) and in forecasting proved inferior to naive time-series models in *ex post* dynamic simulations (Meese and Rogoff, 1983). By the close of the decade the MAER, as also PPP, were largely written off as failures.[4]

Then during the past decade opinion shifted again. The negative results that had coloured thinking about the MAER and PPP had come almost exclusively from studies based upon data for the float alone. In most instances these data spanned somewhere from a decade to a decade and a half. Given the persistence of movements in real exchange rates, it became apparent that these samples were simply too short for reliable statistical inference. Later studies confronted this problem by using substantially expanded data sets, either long historical time series or multi-country panel data for recent decades. The time-series studies typically have applied cointegration techniques to exchange rate data for a small number of currencies over periods of a century or more in length. Two findings have emerged from these studies: (a) real exchange rates in general appear better characterized as mean-reverting than as unit-root processes; (b) this mean reversion takes a good deal of time, estimated half lives of deviations ranging from three to five years.[5] A number of studies using pooled multi-country data for the current float have reached basically the same conclusions.[6]

A factor supporting these latter findings is the change that has occurred in the floating-rate data themselves. First of all, the time span covered by the data relative to the mid-1980s has roughly doubled. Second, and perhaps more important, the volatility that characterized real and

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[4] See Boughton (1988) for a comprehensive review of these models. Note that some studies conducted during these years did, in fact, lend support to the MAER. See, for example Somanath (1985), and for, Canada, both La France and Racette (1985) and Marquez and Schinasi (1988). Later studies for Canada include Backus (1984), Boothe and Poloz (1988), Choudhry *et al.* (1991), Taylor (1996) and Choudhry and Lawler (1997).

[5] The list of such studies is now voluminous. Examples include Abuaf and Jorion (1990), Diebold *et al.* (1991), Lothian and Taylor (1996), Taylor (1996), Sarno and Taylor (1998), and Taylor and Sarno (1998). For historical evidence on the performance of PPP over the past two centuries see Lothian and Taylor (1996). Recent studies of the current float include Flood *et al.* (1995), Lothian (1995), Mark (1995), Sarno and Taylor (1998), and Taylor and Sarno (1988).

[6] See Frankel and Rose (1996) and Lothian (1997) for evidence derived from multicountry panel data for this period. Evans and Lothian (1993), Mark (1995), Sarno and Taylor (1998) and Taylor and Sarne (1998) investigate time series data for a number of the major currencies separately under the float.

nominal exchange rate behaviour in that era – particularly of US dollar rates the causes of which are still being debated – has largely dissipated (Lothian, 1998). The signal in the data has therefore very likely increased both in absolute terms and as a ratio to the 'noise'.

A closely related question is whether a long-term relationship similar to PPP exists between nominal exchange rates and nominal money stocks. Several recent studies have considered this issue. Evidence of the improved performance of MAER models also has begun to accumulate. MacDonald and Taylor (1993, 1994), using cointegration tests and error correction models have presented evidence supportive of the MAER for both the dollar–mark and the dollar–pound exchange rates. Recently, Dutt and Ghosh (2000) found evidence, consistent with the MacDonald and Taylor studies, conducting tests on fixed rate regime (as well as floating rate period) for the Japanese yen–US dollar exchange rate. Other researchers, however, have reported mixed results. Ballie and Pecchenino (1991) conducted separate cointegration tests for the money-demand and PPP relations for the United Kingdom and the United States and although they were able to reject the hypothesis of no cointegration for money demand, they were unable to do so for PPP. DeJong and Husted (1993), using data for five major currencies relative to the US dollar, also presented evidence inconsistent with the MAER.

It is suspected that the reversal of evidence with regard to both PPP and MAER that has taken place over the past half decade can be attributed both to the use of more powerful tests and to more appropriate research design. Earlier studies for the most part used whatever monthly or quarterly data existed up until that point and estimated relatively simple regression models. More recent studies, in contrast, have used substantially expanded data sets – either very long time series or panel data – and or more powerful statistical techniques.

III. DATA AND METHODOLOGY

The data

Monthly data from the *International Financial Statistics* book volumes covering the 1974 to 1993 period was used in this study. The exchange rate used is US dollar per Canadian dollar (line *ag* in monthly IFS data tape). For both countries, the monetary aggregate is M1 (line 34); the income measure is industrial production (line 66C); and the interest-rate measure is the long-term government bond yields (line 61).

Methodology

The cointegration technique pioneered by Engle and Granger (1987), and extended by Johansen (1988), Stock

and Watson (1988) and Johansen and Juselius (1990), among others, has been used as it allows estimation of both long- and short-run relationships without having to difference the data. In testing for the presence of these relationships among the variables included in the data set, the Johansen (1988) procedure is employed. This procedure was chosen because work by Gonzalo (1994) has demonstrated that the Johansen procedure has superior properties to the other methods of testing for cointegration. A brief outline of the procedure follows.

The Johansen procedure requires specifying a kth order vector autoregression (VAR) model for an $n \times 1$ vector of $I(d)$ variables, X_t:

$$X_t = \mu + A_1 X_{t-1} + \cdots + A_k X_{t-k} + \varepsilon_t \qquad (5)$$

where each of the A_i matrices is an $(n \times n)$ matrix of parameters, μ is a deterministic term and ε_t is a vector of residuals which is assumed to be an i.i.d. Gaussian process. This system of equations can be reparameterized in the following error correction form:

$$\Delta X_t = \mu + \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-i} \Delta X_{t-k+i} + \Pi X_{t-k} + \varepsilon_t, \text{(II.)} \qquad (6)$$

where $\Pi = -I + A_1 + \cdots + A_k$, and $\Gamma_i = -I + A_1 + \cdots + A_i$, $i = 1, \ldots, k - 1$. Γ_i represents the matrix of traditional first-difference coefficients that captures the short-run dynamics and Π represents the long-run impact matrix. Important to this study, the variable Π embodies information on the long-run relationships between variables comprising the data set. As such, it is the rank (r) of Π which indicates the number of cointegrating vectors. If Π is of full rank, no cointegration is present as all series are themselves stationary and a conventional regression problem is confronted. On the contrary, if Π has a rank of zero, no stationary long-run relationships are present and Equation 2 is reduced to a VAR in differences. However, if Π has rank $r, 0 < r < n$, it can be factored as $\Pi = \alpha \beta N$, where α and β are $n \times r$ matrices. The r columns of β $\beta N X_t$ is stationary although X_t is not, while $n - r$ represents the number of common stochastic trends. The ith row of α tells us the importance of each of these r cointegrating vectors are to the dynamics of the ith equation. The individual values α_{ir} represent the speed with which the ith series adjusts to the rth cointegrating vector, larger values indicate a more rapid speed of adjustment. Thus not only can the existence of equilibrium relationships be determined, but also the relative speed of adjustments of each cointegrating vector to disequilibrium shocks.

The Johansen procedure uses two likelihood ratio (LR) statistics that test for cointegrating vectors – the 'trace statistic' and the λ_{max} statistic. The former tests the hypothesis that there are at most r distinct cointegrating vectors, while the latter tests the hypothesis of $r + 1$ cointegrating vectors given r cointegrating vectors. These statistics follow non-standard distributions. The asymptotic distributions for

both LR statistics were tabulated and presented in Johansen and Juselius (1990) and Osterwald-Lenum (1992).

Johansen and Juselius (1992) point out that some ambiguity exists in determining the appropriate number of significant eigenvalues. The distributions of the test statistics depend only upon the number of nonstationary components, but, because of differences in the specification of the alternative hypothesis, the critical values associated with the 'λ max' and 'trace' statistics often lead to different conclusions. This dilemma is usually the result of the low power of the test when the cointegrating relationship is quite close to the nonstationary boundary. In the empirical analysis, however, it is found that the results are consistent across both test statistics at the 95% confidence level.

IV. EMPIRICAL RESULTS

First, two tests are performed to investigate the presence of unit roots in each of the series using the Augmented Dickey–Fuller test. The first tests for the presence of a unit root with zero trend in the series, while the second tests for the presence of a unit root with a nonzero linear trend. The results are reported both in levels and first differences in Table 1. In both cases, the null hypothesis cannot be rejected, indicating the presence of a unit root with a trend in each of the series included in the data set.

The presence of a unit root in each series prompted a proceedure with the Johansen's method in testing for the presence of cointegration among the exchange rate and the set of macroeconomic variables. These results are presented in Table 2. Evidence from both the 'λmax' and 'trace' statistics indicates that there are four cointegrating vectors.[7]

The evidence of four cointegrating vectors indicates that there are three common stochastic trends. These results are consistent with an equilibrium subspace between the

Table 2. *Johansen test for cointegration among* s, y, y^F, m, m^F, i, $i^{F a,b}$

Panel A: Eigenvalues in descending order

0.663	0.473	0.401	0.342	0.173	0.0928	0.003

Panel B

	Critical values					
	$k = 4$		λ_{max}		Trace	
Null	8_{max}	Trace	95%	90%	95%	90%
$r = 0$	76.06	206.38	44.06	42.06	124.25	118.99
$r <= 1$	44.79	130.33	39.43	36.35	95.18	90.39
$r <= 2$	35.88	85.54	33.32	30.84	70.60	66.49
$r <= 3$	29.33	49.65	27.14	24.78	48.28	45.23
$r <= 4$	13.32	20.32	21.07	18.90	31.53	28.71
$r <= 5$	6.82	7.00	14.90	12.91	17.95	15.66
$r <= 6$	0.19	0.19	8.18	6.50	8.18	6.50

Notes: [a] Trend maintained.
[b] k refers to the number of lags.

exchange rate and domestic and foreign industrial level of outputs, money supplies and long-term interest rates. Thus, strong support was found for the monetary model of exchange rates. Except for the study of MacDonald and Taylor (1994), this finding contrasts with those of earlier studies of the monetary model of exchange rates and the more recent studies which used the cointegration methodology to test this model. The difference in results is attributed to two factors. First, as argued by MacDonald and Taylor (1994), previous studies used the cointegration approach with single equation methodology as proposed by Engle and Granger (1987). As is well documented in the literature, this regression-based approach is inefficient in that it fails to fully capture the dynamics of the data. Additionally, the cointegration results are sensitive to the

Table 1. *ADF(4) statistics*

	Levels		First differences	
	Without trend[a]	With trend[b]	Without trend[a]	With trend[b]
y^*	−3.222	−3.205	−5.6093	−5.5990
y	−0.455	−3.022	−5.3295	−5.3310
m^*	0.131	−2.354	−4.9018	−5.2863
m	1.209	−1.692	−6.1128	−6.2720
i^*	−1.609	−1.796	−6.9162	−6.9728
i	−1.340	−1.907	−6.2590	−6.3208
s^*	−2.374	−2.190	−7.2128	−7.2300

Notes: [a]95% critical value = −2.901
[b]95% critical value = −3.470

[7] This result is consistent across multiple VAR lag structures. However, four lags produce an error series that was white noise and minimized the AIC criterion.

Table 3. *The estimated cointegrated vectors partitioned into stationary components* ($\exists_i$) *and weights* ($\forall_i$)

Panel A: Eigenvectors

	$\exists_1$	$\exists_2$	$\exists_3$	$\exists_4$
s	-1.000	-1.000	-1.000	-1.000
i	0.423	-0.169	-0.067	0.026
i^*	-0.877	0.229	0.042	0.006
y	0.258	0.001	0.001	-0.029
y^*	0.123	0.023	-0.002	0.009
m	0.008	0.001	0.001	-0.001
m^*	-0.165	-0.010	-0.007	0.014

Panel B: Cointegrating weights

	$\forall_1$	$\forall_2$	$\forall_3$	$\forall_4$
s	-0.003	-0.012	0.205	0.221
i	0.327	1.278	-4.275	-0.072
i^*	0.452	-0.380	-10.956	3.467
y	0.013	-2.587	-1.969	0.699
y^*	0.427	-5.599	22.847	-7.316
m	0.008	0.001	0.001	-0.001
m^*	1.466	-16.909	0.844	-8.314

specific normalization used. Secondly, since those earlier studies, the substantial depreciation in real US dollar exchange rates in the mid-1980s that occurred in following the Plaza and Louvre Accords, and which more or less offset its earlier appreciation, has had sufficient time to be fully reflected in the present set of variables.

In Table 3, the eigenvectors and weights are presented. The weights are the estimated α coefficients and can be interpreted as the average speed of adjustment of each series towards the equilibrium subspace. As such, a large (small) coefficient reflects a high (low) speed of adjustment. The important factor to note about the lower panel of Table 3 is the difference in the speeds of adjustment of the two countries' variables. Specifically, in all cases the Canadian variables are substantially larger than those of the USA, thus indicating that the Canadian variables responded much faster to disequilibrium shocks than do the US variables.

MacDonald and Taylor (1994) pointed out that if the flexible-price monetary approach is correct, then the coefficient on y and y^* should be negative and positive, respectively with numerical values equal to income elasticities from the domestic and foreign demand functions; the coefficients on i and i^* should be positive and negative, respectively, with the size of the coefficients similar to interest rate semi-elasticities obtained from money demand functions.

Finally, the coefficients on m and m^* should be positive one and negative one respectively. The results displayed in Table 3 are in general consistent with these expectations. The first cointegrating vector appears to support the interest-rate relationship, with the industrial-output relationship supported by the fourth cointegrating vector. Note however, that although the second and third cointegrating vectors possess the hypothesized sign for the money supply relationship, the size of the coefficients are significantly different from one.[8]

According to Johansen and Juselius (1992), the eigenvalues can be used as a measure of the relative strength of the long-run relationships, with larger values indicating that the corresponding cointegrating vector(s) are more correlated with the stationary part of the process. Table 2 contains estimates of the eigenvalues in descending order of magnitude with the corresponding cointegrating vectors are displayed in Table 3. These results indicate that the exchange-rate relationship is most correlated with the stationary part of the process, the industrial-output relationship is the least correlated, and the monetary-policy relationship is being the intermediate case. The evidence that the interest-rate relationship is most correlated with the equilibrium subspace, may reflect the fact that the financial markets between these countries are relatively well integrated.[9]

V. CONCLUSIONS

The results reported in this paper indicate that there are four cointegrating relationships defining an equilibrium subspace between the nominal exchange rate and the three pairs of macroeconomic variables characterizing the system of equations. The finding of cointegration among these variables provides strong support for the long-run monetary model of exchange rates. At the same time, it provides indirect evidence in favour of long-run purchasing power parity (PPP) between the US dollar and the Canadian dollar during our sample period.

In contrast to widespread belief, these results suggest, therefore, that the monetary approach to exchange rates, like the PPP relation that is one of its underpinnings, continues to be of use empirically as a long-run predictive tool. It provides an anchor of sorts to the nominal exchange rate – a value from which the nominal exchange rate cannot wander indefinitely, and hence a guide for judging long-run exchange-rate behaviour. These results are consistent with those presented by Taylor (1996) for the Canadian dollar–US dollar, Dutt and Ghosh (2000) for the Japanese yen–US dollar, and MacDonald and Taylor

[8] Hypothesis tests that the coefficients are equal to $+1$ and -1 were not supported.
[9] Similar strong support for the monetary model of exchange rates is evidenced in earlier investigation attempts on long-run exchange rate behaviour of US and Canada dollar (Francis *et al.* 2000).

(1993, 1994) for dollar–mark and dollar–pound exchange rates respectively.

The evidence portrayed in this study is particularly interesting because they are derived from data for two countries in which monetary policies have been remarkably similar. Most previous studies of exchange-rate behaviour – particularly those using long-term time series – have been for countries in which differences in average monetary growth have been truly substantial. In such instances, the possibility arises of this one observation – the difference in means – dominating the statistical findings. The data, however, are largely immune to this problem.

In future work it is hoped to extend the analysis to other industrial-country exchange rates in this and earlier time periods, and also to investigate further the short-run behaviour of these exchange rates.

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