MODELLING DEMAND FOR BROAD MONEY IN AUSTRALIA*

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The existence of a valid long-run money demand function is still important for the conduct of monetary policy. It is argued that previous work on the demand for money in Australia has not been very satisfactory in a number of ways. This paper examines the long- and short-run determinants of the demand for broad money employing the Johansen cointegration technique. Using quarterly data for the period 1976:3–2002:2, this paper finds, *inter alia*, that the demand for broad money is cointegrated with real income, the rate of return on 10-year Treasury bonds, the cash rate and inflation. It appears that a disequilibrium in the demand for money can affect the efficacy of interest rate policy in the long run via its impact on future output growth and output gap.

I. Introduction

Australia's approach to monetary policy has undergone significant changes since 1976. From the mid-1970s until 1985, based on the assumption of a strong and persistent relationship between inflation and the supply of money, monetary policy was conducted by targeting the annual growth of M3. However, in 1985 this policy was abandoned because deregulation of the financial system made M3 a misleading guide to the stance of monetary policy (Grenville 1990). From 1985 to 1989 a "checklist approach" was adopted, whereby a multitude of indicators such as, monetary aggregates, the GDP growth rate, the shape of the yield curve, exchange rates, and the unemployment rate were considered prior to the implementation of monetary policy. The checklist approach was also unsuccessful and finally discontinued in 1989 due to the impossibility of monitoring the above indicators which could provide contradictory policy signals.

Since 1989, the approach taken by the Reserve Bank of Australia (RBA) to monetary policy has been to set the official cash rate in the money market. Following many other OECD countries, inflation targeting has become the ultimate goal of monetary policy in Australia since 1993 (Juttner and Hawtrey 1997). It should be borne in mind that monetary aggregates are still important in this new era of inflation targeting (Hayo 1999). According to Goldfeld (1994) the relation between the demand for money and its main determinants is an important building block in macroeconomic theories and is a crucial component in the implementation of monetary policy.

But why is this so? Is the demand for money more or less important given switches in the emphasis of policy, e.g. from a quantity to a rate based approach? There is no consensus among economists when it comes to these controversial issues. Monetarists believe that a stable demand for money is still a crucial intermediate step to keep the rate of inflation within the

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target range, as in the medium term one would expect real output to grow at its potential or the natural rate. On the other hand, from a Post-Keynesian point of view, the endogeneity of the supply of money renders this question irrelevant, as the RBA cannot control the supply (realised amount) of money in any case. Main stream Keynesian and New Keynesian perspectives would be somewhere in between these two polar positions.

One may argue that the predictability of monetary aggregates will be much less relevant under an interest rate rule as compared to a money supply growth-rule, simply because short-term interest rates have proved to be much less volatile than the money supply growth. In the short run, interest rates can be much more easily and effectively controlled than the money supply (no matter how narrowly defined). This system of controlling the money market rate was not possible under a fixed exchange rate regime. Thus the move to a flexible rate regime has also reduced the concern about the stability of money demand.

However, it can be conjectured that a disequilibrium in the demand for money may affect the efficacy of interest rate policy in the long run via its impact on output gap. This paper provides some empirical support that the real money gap (defined as the difference between the real money stock and the long-term equilibrium real money stock) contains significant information about output gap and the future path of output growth. Therefore, it can be argued that a well-functioning long-run money demand function is still important in this new era of inflation targeting, particularly when the output gap is considered to be a very important determinant of interest rate changes. There are a number of studies that highlight the importance of a well behaved long-run money demand function because the "real money gap" helps to forecast future changes in output gap and/or inflation (see, *inter alia*, Laidler 1999, Gerlach and Svensson 2001, and Siklos and Barton 2001).

Against this background, this paper finds empirical evidence that a disequilibrium in the demand for broad money helps to forecast future changes in real output growth, albeit with no evidence in relation to change in inflation. It is not a valid argument to focus exclusively on a single policy instrument and entirely neglect an important information variable because both the interest rate and monetary aggregates do matter for policy impacts. Given that the output gap is deemed to be an important factor in explaining inflation (as supported by the Taylor rule), one can conclude that the real money gap indirectly affects inflation via its direct influence over the growth of output or output gap. It is essential to track both the cash rate and the money stock in order to assess precisely what monetary policy is doing to the economy. On this same issue Laidler (1999, p. 26) posits that monetary aggregates should not be used "as the only target of monetary policy", but rather as a supplementary intermediate target variable in a regime whose principal anchor is an inflation goal.

A number of studies have already been undertaken to investigate the demand for money in Australia. The review of literature on the demand for money in Australia briefly presented below indicates a growing consensus among economists that broad money (BM) should be considered as an appropriate indicator of monetary aggregate, particularly after the 1980s. However, from 1960 to 1980 there were several studies which reported conflicting results on this issue.

Cohen and Norton (1969) can be considered as pioneers of money demand analysis in Australia. Using a modified stock adjustment model, they estimate several money demand functions with the limited available quarterly data for various monetary aggregates in Australia. Unlike Adams and Porter (1976) who argue against stability of M1, Pagan and Volker (1981) employ a conventional specification of the demand for money function and found a stable relationship for M1. Sharpe and Volker (1977) and Lim and Martin (1991) in their study of M3 in Australia argued for the stability of the money demand function, while Blundell-Wignall and Thorp (1987), and Orden and Fisher (1993) modelled M1, M3, and BM, and found exactly the opposite. de

Brouwer, Ng and Subbaraman (1993) and Juselius and Hargreaves (1992) use Australian data and correctly conclude that the number of cointegrating vectors and their stability are very sensitive to the choice of scale variable, e.g. GDP or GNE (gross national expenditure), and the measure of money. Using the Johansen test, de Brouwer, Ng and Subbaraman (1993) have also examined various measures of money, different interest rates, and scale variables, and concluded that there is evidence of cointegration between money, income and the interest rate, particularly for BM.

Using the Engle and Granger two-step methodology and quarterly data for the 1970:1–1993:1 period, Hoque and Al-Mutairi (1996) find a long-run relationship between narrow money, output, the interest rate, and price level. They conclude that this long-term relationship shows no sign of instability in the face of financial innovation and deregulation in the 1980s. It can be argued that the model formulated by Hoque and Al-Mutairi is misspecified as it includes only one interest rate (the two-year Treasury bill rate) in the equation for money demand, ignoring the process of financial asset substitution.

Felmingham and Zhang (2001) have identified this misspecification, and included a more appropriate measure of opportunity cost of holding money (i.e. the interest rate spread defined as the difference between short- and long-run interest rates) in their cointegrating vector of the demand for BM. They employ the Johansen cointegration technique and find that there exists a cointegrating vector linking BM with real GDP, the interest rate spread and inflation over the period 1976:3–1998:4. They have also performed several residual-based tests for cointegration to identify a structural break in their long-run relationship for BM. Felmingham and Zhang (2001) conclude that this long-run relationship was subject to regime shifts in 1991. They (p. 50) mistakenly deflate nominal GDP with the consumer price index (CPI) and also assume that the semi-elasticities of the interest rate on money itself and the interest rate outside money have equal magnitude but with opposite signs without testing such an important restriction. Furthermore, they have not estimated a short-run dynamic model for BM and no weak exogeneity testing was undertaken either.

This paper updates the sample and addresses the problems and shortcomings associated with the previous work on the demand for BM. The structure of the paper is as follows. In Section II a theoretical model is postulated which captures the long-run demand for money using the Johansen multivariate cointegration technique. Definitions of the variables, sources of the quarterly data employed as well as the unit-root results using the Augmented Dickey-Fuller (ADF) and the Kwiatskowski-Phillips-Schmidt-Shin (KPSS) tests are presented in Section III. The empirical econometric results for the long- and short-run demand for money, as well as policy implications of the study are also discussed in this section. Section IV provides some concluding remarks.

II. Theoretical Framework

Conventionally the demand for money is specified as a function of real income, a long-run interest rate on substitutable non-money financial assets, a short-run rate of interest on money itself, and the inflation rate. More specifically, following, *inter alia*, Ericsson (1998), Beyer (1998) and Coenen and Vega (2001), the demand for money function is specified as follows:

$$m_t - p_t = \gamma_0 + \gamma_1 y_t + \gamma_2 RL_t + \gamma_3 RS_t + \gamma_4 \Delta p_t + \varepsilon_t \qquad (1)$$

where m is nominal money demanded (as an observable variable) , p is the price level, y is a scale variable, RL is the long-run rate of return on assets outside of money and RS is the short-run rate of interest on money itself. All variables shown in lowercase (i.e. m, y, and p) are in logs and the remaining variables (i.e. RL and RS) are in levels. As a result, γ_1 denotes the income elasticity of the demand for money, whereas γ_2, γ_3, and γ_4 are semi-elasticities of RL, RS, and the inflation rate with respect to money demand, respectively. In practice if γ_2 and γ_3 have coefficients of equal magnitude but opposite signs, equation (1) can also be rewritten in the following form:

$$m_t - p_t = \gamma_0 + \gamma_1 y_t + \gamma_2(RL_t - RS_t) + \gamma_4 \Delta p_t + \varepsilon_t \tag{2}$$

Following earlier studies mentioned above, in order to avoid dealing with I(2) variables (m and p), equations (1) and (2) are usually employed to model real money balances, supporting the price homogeneity assumption. The expected sign and magnitude of the coefficient for the scale variable is as follows: if $\gamma_1 = 1$, the quantity theory applies; if $\gamma_1 = 0.5$, the Baumol-Tobin inventory-theoretic approach is applicable; and if $\gamma_1 > 1$, money can be considered as a luxury or it can also be interpreted as an indication of neglected wealth effects. According to Ball (2001), an income elasticity of less than unity has a number of implications for monetary policy. For instance, one may conclude that the Friedman rule is not optimal in this case and the supply of money should grow more sluggishly than output to achieve the goal of price stability (Ball 2001, p. 36). For a detailed discussion of controversy about the quantity theory see Laidler (1991).

It is also expected that RL, as a proxy for the yields on outstanding Treasury bonds, has a negative sign or $\gamma_2 < 0$, whereas the coefficient for the short-run rate of interest is positively correlated with money demand, or $\gamma_3 > 0$. The annualised rate of inflation $\Delta p = \Delta_4 p_t = \ln(P_t) - \ln(P_{t-4})$, is considered as a proxy to measure the return on holdings of goods and its coefficient should thus be negative, i.e. $\gamma_4 < 0$, as goods are an alternative to money. The exclusion or inclusion of inflation in this equation is an issue of dynamic specification. In the literature the inclusion or exclusion of Δp in equation (1) is the subject of some ongoing controversy. In fact, the exclusion of inflation imposes equality of the short- and long-term elasticities of nominal BM with respect to prices. Many previous works such as Lim and Martin (1991), Bärdsen (1992), Hendry and Ericsson (1991), and Wolters, Teräsvirta, and Lütkepohl (1998) have included the rate of inflation (actual) in the demand for money function and have strongly rejected this restriction.

More specifically, the inclusion of the annualised rate of inflation in dynamic models of money demand is of particular importance for three reasons. First, it allows us to reparameterise the money demand models in terms of real money balances and inflation. According to Coenen and Vega (2001, p. 729) "such reparameterisation allows for the theoretically plausible hypothesis of long-run price homogeneity of money demand but does not impose any untested (and frequently empirically rejected) common factor restriction of short-run price homogeneity". When m and p are I(2) and $m - p$ is I(1), as it is the case in this study, one can resort to this reparameterisation to map an I(2) system into an I(1) system.

Second, previous studies (e.g. Ericsson 1999) highlight the importance of the inclusion of the actual rate of inflation as an important determinant of constant-parameter empirical models of money demand. This argument is justified on the basis of the fact that inflation is a good proxy for the opportunity cost of holding money rather than real assets. Third, it is argued that the inclusion or exclusion of Δp in models of real money demand is a dynamic specification issue, which should be subject to an empirical testing. For a comprehensive discussion of the literature on money demand see also, *inter alia*, Laidler (1993) and Hoffman and Rasche (2001).

If empirical results do not reject the null hypothesis of $\gamma_1 = 1$, then the (inverse) long-run velocity of BM can be obtained by:

$$(m_t - p_t - y_t) = \gamma_0 + \gamma_2 RL_t + \gamma_3 RS_t + \gamma_4 \Delta p_t + \varepsilon_t \quad (3)$$

In order to have a valid model for the money demand function, there should be at least one cointegrating vector in the system. The Johansen (1991, 1995) multivariate cointegration technique is used in this paper to test the existence of a long-run equilibrium relationship among the variables specified in equation (1). A brief description of this technique is presented below.

Let us consider the following VAR of order q:

$$y_t = A_1 y_{t-1} + A_2 y_{t-2} + \ldots + A_q y_{t-q} + \varepsilon_t \quad (4)$$

where y_t is a k-vector of I(1) variables (e.g. in this study $k = 5$ and the variables are $m - p$, y, RL, RS, and Δp), and ε_t is a vector of white noise residuals. Following Johansen (1991, 1995) equation (4) can also be rewritten as:

$$\Delta y_t = \Pi y_{t-1} + \sum_{i-1}^{q-1} \Gamma_i \Delta y_{t-i} + \varepsilon_t \quad (5)$$

where $\Pi = \sum_{i=1}^{q} A_i - I$, and $\Gamma_i = -\sum_{j=i+1}^{q} A_j$.

The rank (r) of Π determines the number of cointegrating vectors. If Π has a reduced rank (i.e. $r < k$), then there exist $k \times r$ matrices α and β each with rank r, where $\Pi = \alpha\beta'$ and $\beta' y_t$ is stationary. The elements of α represent the adjustment parameters and each column of β in the literature is referred to as the cointegrating vector. Thus the important issue is how to determine the number cointegrating vectors (or r). In this paper both the trace statistics and the maximum eigenvalue statistics will determine r. The trace statistics test the null hypothesis of r cointegrating relations against the alternative of k cointegrating equations. On the other hand, the maximum eigenvalue statistics test the null of r cointegrating vectors versus the alternative of $r + 1$ cointegrating relations. For more details see Johansen (1991, 1995).

An important step before using the Johansen multivariate technique is to determine the time series properties of the data. This is an important issue since the use of non-stationary data in the absence of cointegration can result in spurious regression results. To this end, two unit root tests, i.e. the ADF test, and the Kwiatskowski-Phillips-Schmidt-Shin (KPSS) (Kwiatkowski *et al.* 1992) test, have been adopted to examine the stationarity, or otherwise, of the time series data. In this paper the lowest value of the Akaike Information Criterion (AIC) has been used as a guide to determine the optimal lag length in the ADF regression. These lags augment the ADF regression to ensure that the error term is white noise and free of serial correlation. Unlike the ADF test, the KPSS test has the null of stationarity, and the alternative indicates the existence of a unit root.

III. Empirical Results and Policy Implications

The nominal demand for BM rose substantially from $49.2 billion in the third quarter of 1976 to $546.5 billion in the second quarter of 2002, an average growth of 2.3 per cent per quarter or 9.2 per cent per annum. What are the major long- and short-run determinants of the demand for BM during the last four decades? Based on the theoretical framework discussed in Section II, the objective of this paper is to answer this question.

Before embarking on our empirical quest, it is important to look at the sources and definitions of the data presented in Table I. Quarterly time series data employed for the period 1976:3–2002:2 are as follows: nominal broad money (m), the consumer price index (p), real GNE (y), ($m - p$), the rate of return on 10-year Treasury bonds as a proxy for RL, and the official cash rate as a proxy for RS. The three variables of m, p, and y are seasonally adjusted. Following

Table I Sources and definitions of the data employed

Variables	*Unit*	*Sources*
Broad money or *M* and $m = \ln(M)$	\$ million and seasonally adjusted (sa)	RBA (2002), tables D03 and F01.
The cash rate or *RS*	Fraction	
The rate of return on 10-year treasury bond or *RL*	Fraction	ABS (2002a), table 31.
The consumer price index or *P* and $p = \ln(P)$	1989–1990 = 100	ABS (2002b)
Real GNE or *Y* and $y = \ln(Y)$	\$ million, sa Chain volume measures, 1999 prices.	ABS (2002c), table 5.

the literature, *RL*, *RS* and the rate of inflation are expressed as fractions, whereas, the other variables are in logs and thus shown in lowercase. According to de Brouwer, Ng and Subbaraman (1993, p. 10), BM encompasses "M3 plus borrowings from the private sector by non-bank financial intermediaries (NBFIs), less their holdings of currency and bank deposits". They also argue that compared with other measures of money, the evidence of cointegration is stronger when BM is modelled as it: a) is less distorted by financial deregulation and innovations; and b) has a more reliable relationship with GNE. Following the literature, in this paper BM is preferred to other narrower measures of money such as M1 and M3 which can be substantially affected by asset substitution and are also more volatile (Felmingham and Zhang 2001). de Brouwer, Ng and Subbaraman (1993, p. 10), believe that "selection of the income and interest rate variables is largely an empirical matter."

However, the choice of interest rates depends on the measure of money being modelled. While Felmingham and Zhang (2001) considered the weighted average 5- and 10-year Treasury bond interest rates as a proxy for *RL*, and the weighted average of interest rate charged by AMMD (authorised money market dealers), as a proxy for *RS*, this study uses the cash rate and the interest rate on the 10-year Treasury bonds as proxies for *RS* and *RL*, respectively. The rationale for this decision is twofold: the first reason relates to the issue of data reliability, and the second pertains to the nature and recognition of the cash rate as a policy variable in practice. These two reasons are discussed below in more detail.

First, as seen from Figure 1, the cash rate and the AMMD rate move very closely to each other particularly after 1983 (when the exchange rate was floated). However, the data on the cash rate seem more reliable as the average AMMD rate only represented "authorised dealers" before the deregulation and thus it did not cover the interest rates paid by NBFIs, which played a very important role in Australia's financial system over this period. Therefore, the use of the AMMD rate may create a measurement error in the proxy for *RS*.

The second reason pertains to the purpose of this study. The motivation for this study is to estimate the short- and long-run impact of (say) a one per cent change in a policy variable which can be controlled and changed directly by the RBA as a policy variable. This variable is the cash rate and the RBA is the only agent which has exclusive right to set and fine tune it. The official cash rate is considered a good proxy for *RS* as it exerts a great influence on all other interest rates in the money market, e.g. those of Treasury notes and 90-day bank bills.

Ericsson (1998) suggests that long-run rates should not be included in the demand equation for M1. However, if a broader definition of money is modelled, it is essential to incorporate longer-term interest rates in the demand for money function so as to capture financial asset substitutions. This paper examines the demand for "broad" money, and as a result *RL* is best proxied by a "long-run rate" such as the rate of interest on 10-year Treasury bonds, a security with the longest maturity for which the quarterly time series data are available. The broader the

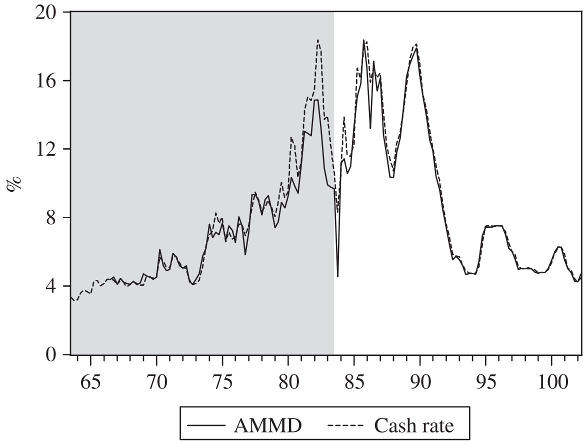

Figure 1. Plot of the AMMD rate and the official cash
Source: Table I.

Table II ADF test results

Variable	*C (constant) and T (trend) in the ADF equation*	*ADF statistics*	*Optimum lag length Using the AIC*
m	*C & T*	−1.60	4
Δm	*C*	−2.38	3
$\Delta\Delta m$	*C*	−5.93*	4
p	*C & T*	−1.50	5
Δp	*C*	−2.50	2
$\Delta\Delta p$	*C*	−9.07*	2
$(m - p)$	*C & T*	−2.48	3
$\Delta(m - p)$	*C*	−7.95*	0
y	*C & T*	−2.60	1
Δy	*C*	−7.74*	0
RL	*C & T*	−2.45	3
ΔRL	*C*	−8.68*	0
RS	*C & T*	−2.99	4
ΔRS	*C*	−5.23*	5

Note: * indicates that, based on the MacKinnon critical values, the corresponding null hypothesis is rejected at the 1% significance level.

definition of money, the longer rates would be more relevant. Besides, the use of a weighted average of 5- and 10-year Treasury bonds, as employed by Felmingham and Zhang (2001), may be susceptible to measurement errors associated with likely inaccuracy of "true weights" in the computation of such a measure.

Prior to undertaking an empirical investigation of the sources of demand for BM, it is essential to determine the time series properties of the data. In order to make robust conclusions about stationarity or otherwise of the data, the ADF and the KPSS tests are utilised. The empirical results of the ADF tests are summarised in Table II. According to the results of the ADF test, both m and p are I(2), whereas $m - p$ is I(1), indicating that p and m become stationary after second differencing, whereas $m - p$ reaches stationarity after first differencing. This is the main reason

Table III KPSS statistics for null of level and trend stationarity

Variables	*C (constant) and T (trend) in the KPSS equation*	*Lag truncation parameter (l)*								
		0	*1*	*2*	*3*	*4*	*5*	*6*	*7*	*8*
m	*C & T*	2.39*	1.211*	0.817*	0.620*	0.502*	0.424*	0.368*	0.327*	0.295*
Δm	*C*	3.06*	1.90*	1.42*	1.13*	0.963*	0.846*	0.756*	0.688*	0.636*
$\Delta\Delta m$	*C*	0.031	0.044	0.061	0.062	0.069	0.082	0.088	0.088	0.090
p	*C & T*	2.552*	1.300*	0.878*	0.667*	0.541*	0.547*	0.397*	0.352*	0.318*
Δp	*C*	4.778*	3.085*	2.278*	1.819*	1.522*	1.327*	1.182*	1.075*	0.986*
$\Delta\Delta p$	*C*	0.052	0.094	0.117	0.135	0.133	0.147	0.149	0.170	0.170
$(m - p)$	*C & T*	0.784*	0.402*	0.275*	0.212*	0.174*	0.150*	0.133	0.120	0.111
$\Delta(m - p)$	*C*	0.189	0.155	0.136	0.120	0.111	0.107	0.104	0.102	0.101
y	*C & T*	0.900*	0.472*	0.330*	0.261*	0.221*	0.194*	0.176*	0.163*	0.153*
Δy	*C*	0.133	0.107	0.098	0.093	0.093	0.098	0.102	0.108	0.116
RL	*C & T*	1.641*	0.843*	0.576*	0.443*	0.364*	0.312*	0.276*	0.249*	0.228*
ΔRL	*C*	0.140	0.122	0.119	0.110	0.108	0.107	0.107	0.106	0.107
RS	*C & T*	1.407*	0.731*	0.506*	0.393*	0.326*	0.216*	0.252*	0.230*	0.214*
ΔRS	*C*	0.185	0.155	0.154	0.148	0.135	0.129	0.136	0.141	0.147

Note: * indicates that, based on the Kwiatkowski-Phillips-Schmidt-Shin (1992) critical values, the corresponding null hypothesis of stationarity is rejected at the 5% significance level. The 5% critical values are: 0.146 (when both *T* & *C* are included in the KPSS test) and 0.463 (when only *C* is included in the KPSS test equation).

why in many studies $(m - p)$, instead of m, is modelled in equation 1. All the other variables, i.e. (y, RL, RS, and Δp), are I(1).

Table III presents the results of the KPSS test for level (with constant only) and trend stationarity (with both a constant and trend) up to a maximum of eight truncation lags (l). As seen from Table III, irrespective of the number of truncation lags and consistent with the ADF test results, m and p are again I(2) and y, RL and RS are I(1). It should be noted that, according to the KPSS test results, the variable $(m - p)$ is I(1) using a lag truncation parameter of up to five but the addition of more lags results in the reversal of this conclusion. In other words, $(m - p)$ is I(1) if one considers the KPSS statistic from lag zero to lag five but with the use of six to eight lags the KPSS test fails to reject stationarity of this variable. Given the fact that in most cases the problem of serial correlation for quarterly time series data is likely to be of order one to four, a maximum upper bound of four truncation lags (l) will be enough to ensure that autocorrelation is corrected in the KPSS test. Therefore it is assumed that $(m - p)$ is also I(1). In sum, the ADF and KPSS tests for unit roots support the view that m and p are I(2), and the remaining variables, which will be used in equation (1), *viz.*, $(m - p)$, y, RL, RS, and Δp, are I(1) for the sample under investigation.

Since all the variables in equation 1 are I(1), the Johansen (1991, 1995) multivariate cointegration technique can now be used to test the existence of a long-run equilibrium relationship for BM. In addition to the five variables discussed earlier, a dummy variable (du) has been considered. This intercept dummy variable is equal to one before 1983 and otherwise zero. The inclusion of this dummy variable (du) related to the Australian currency being floated in 1983, producing an important effect on the financial and monetary system. It is assumed that this dummy variable affects the vector error correction (VEC) model but not the cointegrating vector(s). Following Coenen and Vega (2001), an unrestricted intercept and a linear trend in the variables but not in the cointegrating vectors enter the system. The first important step in this test is to determine the optimal lag length (q) in equations (4) or (5). Allowing for an upper

Table IV Johansen test for cointegration

Hypothesised No. of CE(s)	*Eigenvalue*	*Trace statistic*	*1% critical value*	*Max. Eigenvalue statistic*	*1% critical value*
None	0.323	89.7*	76.1	39.4*	38.8
At most 1	0.182	50.3	54.5	20.4	32.2
At most 2	0.151	29.9	35.7	16.5	25.5
At most 3	0.123	13.4	20.0	13.2	18.6
At most 4	0.001	0.12	6.7	0.124	6.7

Note: * indicates that the corresponding null hypothesis is rejected at 1% significance level.

Table V Standardised cointegrating vector and the corresponding adjustment coefficients

Cointegrating eq	*β coefficients*	*t ratio*	*VEC equation*	*α coefficients*	*t ratio*
			Non-restricted (m − p) model		
$(m-p)_{t-1}$	1	–	$\Delta(m-p)_{t-1}$	−0.153	−2.5
y_{t-1}	−1.100	−32.0	Δy_{t-1}	0.148	2.0
RL_{t-1}	2.010	5.40	ΔRL_{t-1}	−0.035	−1.0
RS_{t-1}	−1.403	−7.2	ΔRS_{t-1}	0.111	1.6
Δp_{t-1}	0.327	2.1	$\Delta\Delta p_{t-1}$	−0.065	−1.1
Constant	4.974	–			
			Restricted (m − p − y) model		
$(m-p-y)_{t-1}$	1	–	$\Delta(m-p-y)_{t-1}$	−0.132	−2.8
RL_{t-1}	3.65	8.2	ΔRL_{t-1}	−0.038	−1.6
RS_{t-1}	−2.38	−7.6	ΔRS_{t-1}	0.078	1.6
Δp_{t-1}	0.45	1.7	$\Delta\Delta p_{t-1}$	−0.049	−1.3
Constant	3.71	–			

band of four lags, three lag selection criteria of the FPE (final prediction error), the sequential modified LR (likelihood ratio) test statistic and the AIC have been employed to determine q. Based on these criteria (not reported here but available from the author upon request), the optimum lag length is $q = 2$. There are a number of other recent studies modelling the quarterly demand for money that have also used an optimal lag length of two. See, for example, Beyer (1998), Coenen and Vega (2001), and Schmidt (2001). Various diagnostic tests indicate that the system of equations with two lags is well-behaved.

However it should be noted that the rank (r) of Π in this study is not sensitive to the lag length. Both the trace and max-eigenvalue tests, using a variety of lags ranging from one to five in separate VAR models, reject a zero cointegrating vector in favour of one cointegrating vector at the one per cent significance level. Table IV reports the results of the Johansen multivariate cointegration test on the demand for BM as formulated in equation (1). As seen there is robust evidence of one cointegrating vector at the one per cent level. Due to space limitations, the cointegration test results using other lags (i.e. 1, 3, 4, and 5 lags) are not reported here but are available from the author on request.

From Table V the long-run parameters are seen to be of consistent sign and orders of magnitude and highly significant. It should be noted that the eigenvalue associated with the first vector (0.323) in Table IV is considerably higher than those corresponding to the other vectors, thereby validating that there exists a unique cointegrating vector in the system. As can be seen from the results obtained from the unrestricted cointegrating vector in Table V, the long-run demand for

Table VI Testing for restrictions on the αs and the βs

The null hypothesis	*Statistic*	*Probability*
$\alpha_{m-p} = 0$	$\chi^2(1) = 4.08$*	0.04
$\alpha_y = 0$	$\chi^2(1) = 2.85$	0.09
$\alpha_{RL} = 0$	$\chi^2(1) = 1.02$	0.31
$\alpha_{RS} = 0$	$\chi^2(1) = 2.37$	0.12
$\alpha_{\Delta p} = 0$	$\chi^2(1) = 0.85$	0.36
$\alpha_y = \alpha_{RL} = \alpha_{RS} = \alpha_{\Delta p} = 0$	$\chi^2(4) = 10.04$*	0.04
$\alpha_{RL} = \alpha_{RS} = \alpha_{\Delta p} = 0$	$\chi^2(3) = 5.8$	0.12
$\gamma_1 = 1$	$\chi^2(1) = 3.84$	0.06
$\gamma_2 = -\gamma_3$	$\chi^2(1) = 5.1$*	0.02

Note: * indicates that the corresponding null hypothesis is rejected at 5% significance level.

BM (after rearranging the vector) is positively related to the own-rate (*RS*) and negatively to both the rate of return on other substitutable financial assets (*RL*) and the annualised rate of inflation.

Table V also shows the estimated adjusted coefficients (αs) which can be used to test for weak exogeneity. The adjustment coefficients contain weights with which cointegrating vector(s) enter short-run dynamics. Given that this study finds only one cointegrating vector, Table V presents the first column of the α matrix. These coefficients measure the speed of the short-run response to disequilibrium occurring in the system. Before proceeding any further, it is essential to test for weak exogeneity of the four variables on the right hand side of equation (1) with respect to $(m - p)$. The Johansen method enables analysts to test for weak exogeneity by imposing zero restrictions on the weighting coefficients of α_y, α_{RL}, α_{RS}, and $\alpha_{\Delta p}$. One should note that the *ec* term is significant and correctly signed (−0.153) in the VEC equation for $(m - p)$.

Table VI, *inter alia*, presents the test results for separate and joint restrictions on the weighting coefficients. As can be seen, using separate zero restrictions on the corresponding αs, the *ec* term, while highly significant for $(m - p)$, is not significant in the short-run dynamic equations for *y*, *RL*, *RS*, and Δp. The weak exogeneity test, by imposing the joint zero restriction of $\alpha_{RL} = \alpha_{RS} = \alpha_{\Delta p} = 0$, reveals that the null cannot be rejected at the one per cent level as $\chi^2(3) = 5.8$ [probability = 0.12]. However, the joint restriction of $\alpha_y = \alpha_{RL} = \alpha_{RS} = \alpha_{\Delta p} = 0$ can be rejected at four per cent. See Table VI.

On the basis of imposing separate and joint restrictions on the adjustment coefficients one can conclude that while three variables of *RL*, *RS* and Δp are weakly exogenous with respect to $(m - p)$ but *y* is not. Thus, it can be stated that *RL*, *RS* and Δp fail to respond to past vector disequilibria or the real money gap. However, because the null of $\alpha_y = 0$ is rejected at the ten per cent level, this means that the real money gap (proxied by ec_{t-1}) has some significant explanatory power for future output growth. As can be seen from the upper section of Table V, the magnitude of the estimated coefficient for α_y indicates that a ten per cent deviation in the demand for broad money from its long-run trend can increase the real growth of output by about 1.5 per cent in each quarter.

As mentioned earlier, a well functioning money demand is important because the deviation of the demand for BM from its long-run path provides some predictive power in explaining or forecasting short-run movements in the output gap and/or inflation. Similar results were found by Laidler (1999), Gerlach and Svensson (2001), and Siklos and Barton (2001). In order to measure the effects of the real money gap (where $ec_t = RMG_t$) on inflation and the output gap in a more direct manner, as examined by Gerlach and Svensson (2001), the following two equations have been estimated:

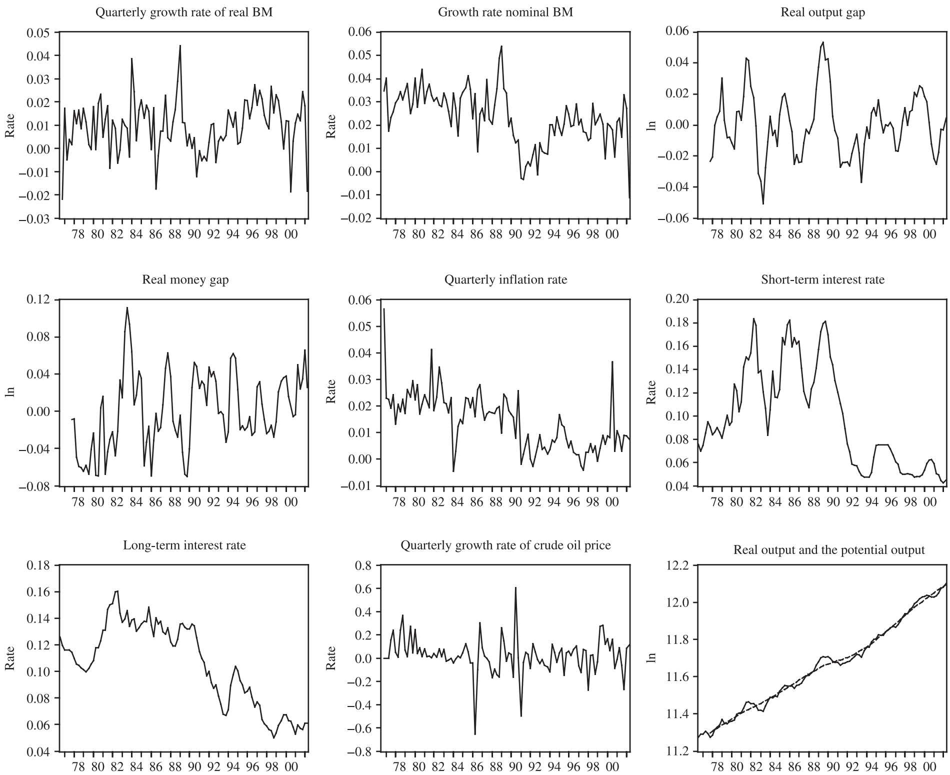

Figure 2. Plots of the data employed-1976:3–2002:2

Source: (1) Table I, (2) ABS (2002a, Table II), and (3) author's calculations.

$$OG_t = \alpha_1 + \sum_{j=0}^{k=4} \lambda_{1j} RMG_{t-j} + \sum_{j=0}^{k=4} \lambda_{2j} \Delta \log(M/P)_{t-j} + \sum_{j=1}^{k=4} \lambda_{3j} (OG)_{t-j} + \varepsilon_{1t} \tag{6}$$

$$\begin{aligned} \Delta \log(P)_t = \alpha_2 &+ \sum_{j=0}^{k=4} \lambda_{4j} RMG_{t-j} + \sum_{j=0}^{k=4} \lambda_{5j} (OG)_{t-j} + \sum_{j=0}^{k=4} \lambda_{6j} \Delta \log(M)_{t-j} \\ &+ \sum_{j=0}^{k=4} \lambda_{7j} \Delta \log(Poil)_{t-j} + \sum_{j=1}^{k=4} \lambda_{8j} \Delta \log(P)_{t-j} + \varepsilon_{2t} \end{aligned} \tag{7}$$

where *OG* denotes an estimate of the output gap (defined as the difference between log(Y) and its long-term trend component using the Hodrick-Prescott filter), *Poil* is the crude oil price index (Australian Bureau of Statistics 2002a, Table 11) and as before *Y*, *M* and *P* are real output, nominal broad money and the CPI, respectively. Figure 2 presents the plots of the data employed.

The OLS and SUR (seemingly unrelated regression) methods have been used to estimate the above two equations over the sample period (1978:4–2002:2). In order to obtain a more parsimonious model, the general to specific method has also been adopted to exclude the insignificant λ_{ij} (where $i = 1, 2, \ldots, 8$) coefficients from the equations. Although it is not the main objective of this paper to model inflation or the real output gap, the empirical results, presented in Table VII, show the extent to which a dis-equilibrium in the money market can affect inflation

Table VII Major determinants of the real output gap and inflation, 1978:4–2002:2

	Dependent variables			
	Output gap: OG_t		*Inflation: $\Delta\log(P)_t$*	
Independent Variables	*OLS Coefficient [t-probability]*	*SUR Coefficient [t-probability]*	*OLS Coefficient [t-probability]*	*SUR Coefficient [t-probability]*
Intercept	−0.0062 [0.0001]	−0.0063 [0.0000]	0.0007 [0.7104]	0.0007 [0.6985]
RMG_t	−0.1110 [0.0017]	−0.1109 [0.0010]	–	–
RMG_{t-1}	0.0916 [0.0101]	0.0921 [0.0066]	–	–
$\Delta\log(M/P)_t$	0.3022 [0.0013]	0.3086 [0.0005]	–	–
$\Delta\log(M/P)_{t-2}$	0.3504 [0.0004]	0.3494 [0.0002]	–	–
OG_{t-1}	0.9380 [0.0000]	0.9368 [0.0000]	0.0708 [0.0747]	0.0708 [0.0656]
OG_{t-2}	−0.2234 [0.0129]	−0.2211 [0.0099]	–	–
$\Delta\log(M)_{t-2}$	–	–	0.1768 [0.0334]	0.1772 [0.0277]
$\Delta\log(Poil)_{t-1}$	–	–	0.0083 [0.0738]	0.0082 [0.0683]
$\Delta\log(P)_{t-2}$	–	–	0.3139 [0.0011]	0.3142 [0.0007]
$\Delta\log(P)_{t-3}$	–	–	0.2939 [0.0021]	0.2927 [0.0016]
$\bar{R}^2$	0.796	0.796	0.523	0.523
DW	1.91	1.90	1.65	1.65

Diagnostic tests:	*Test statistic*	*Prob.*	*Test statistic*	*Prob.*
AR 1–5:	$F(5,83) = 0.380$	[0.8612]	$F(5,84) = 1.284$	[0.2785]
ARCH 1–4	$F(4,80) = 0.299$	[0.8775]	$F(4,81) = 0.3942$	[0.8122]
Normality	$\chi^2(2) = 2.314$	[0.3144]	$\chi^2(2) = 14.383$	[0.0008]**
White heteroskedasticity:				
no-cross terms	$F(12,75) = 1.522$	[0.1350]	$F(10,78) = 0.545$	[0.8531]
cross-terms	$F(27,60) = 1.189$	[0.2836]	$F(20,68) = 0.298$	[0.9982]
RESET	$F(1,87) = 1.760$	[0.1882]	$F(1,88) = 1.742$	[0.1904]

or the real output gap through time and in a dynamic manner. The results clearly indicate that the real money gap (*RMG*) has very significant effects (λ_{10} and λ_{11}) on the output gap. In other words, if money demand fluctuates wildly, the future output gap will vary accordingly. However, *RMG* and its various lagged values were not significant at any conventional level and thus they have been excluded from Table VII. More specifically, the null hypothesis of H_0: $\lambda_{40} = \lambda_{41} = \lambda_{42} = \lambda_{43} = \lambda_{44} = 0$ was examined using the Wald test. Given that $F(5,84) = 0.459$ [P-value = 0.8068], the null could not be rejected at one, five or ten per cent significance levels. Therefore, it is concluded that the dis-equilibrium in the money market may not directly affect inflation. However, as seen from Table VII, the lagged values of output gap, changes in crude oil prices and the growth of nominal broad money were all statistically significant in the inflation equation.

Although the dis-equilibrium in the money market may not determine inflation directly, it has significant effects on the output gap which, in turn, can affect inflation, suggesting an indirect or recursive effect of *RMG* on inflation. These results reinforce our previous claims in that both monetary aggregates and the demand for money should still be considered as useful information variables in this era of inflation targeting. Given the difficulty of forecasting the output gap in Australia this would seem to be a very useful finding.

Attention is now directed to the specification of a short-term demand for money. A number of studies have modelled $\Delta(m - p - y)_t$ rather than $\Delta(m - p)_t$, imposing equality of the long- and short-run income elasticities, e.g. Hayo (2000) in the case of the demand for money in Austria. If the null of $\gamma_1 = 1$ is not rejected, then one can model short-run dynamics of $\Delta(m - p - y)_t$ instead of $\Delta(m - p)_t$.

As seen from Table V, the estimated long-run income elasticity (1.10) is reasonably close to unity which is consistent with the quantity theory of money and other studies for developed countries, e.g. Beyer (1998) in his study of M3 in Germany, Coenen and Vega (2001) in their recent study of M3 in the Euro area, and Ericsson (1998) in his analysis of the narrow demand for money in the UK. Nevertheless, one needs to test formally the $\gamma_1 = 1$ assumption on the cointegrating vector. Table VI also presents the LR test result for this restriction. Given that $\chi^2(1) = 3.84$ [probability = 0.06], one cannot "marginally" reject the null of $\gamma_1 = 1$ at five per cent level.

Attention is now placed on a restricted version of the cointegrating vector. This restricted cointegrating vector links $(m - p - y)$ with RL, RS and Δp, implying that the long- and short-run income elasticities are equal to one. The cointegration results under the assumption of $\gamma_1 = 1$ for this restricted model are shown in the lower part of Table V. The adjustment coefficient for the dependent variable has changed slightly from −0.15 in the non-restricted vector to −0.13 in the restricted cointegrating vector. This coefficient is highly significant, correctly signed and within an acceptable range. For example, Coenen and Vega (2001, p. 737) in their study of M3 in the Euro area and Beyer (1998, p. 60) in his study of M3 in Germany found the corresponding adjustment coefficient to be −0.132 and −0.141, respectively. The restricted cointegrating vector is also presented below.

$$(m_t - p_t) = y_t - 3.7 - 3.65RL_t + 2.38RS_t - 0.45\Delta p_t \tag{8}$$

At this stage one may also want to test if $\gamma_2 = -\gamma_3$. Using the cointegrating vector in equation (8), this restriction has been tested and the result from the LR test is presented in Table VI. Given that $\chi^2(1) = 5.1$ [probability = 0.02], the null hypothesis of $\gamma_2 = -\gamma_3$ is rejected at the two per cent significance level, indicating that RL and RS do not have coefficients of equal magnitude but opposite signs. Thus equation (8) is used to analyse the long-term determinants of the demand for BM. One should note that the most recent study undertaken by Felmingham and Zhang (2001) on the demand for BM in Australia has not tested this hypothesis and assumed that $\gamma_2 = -\gamma_3$.

As seen from equation (8), consistent with the quantity theory of money supporting a long-run income elasticity of unity, a one per cent increase in real income stimulates the real demand for BM by one per cent. Given that the estimated coefficients of −3.65, +2.38 and −0.45 are the semi-elasticities for RL, RS and Δp, respectively, one can convert them to elasticities by multiplying each one of them by the value of its corresponding variable in each quarter. Thus the magnitudes of the resulting elasticities vary depending on the value taken by these variables. For instance, given that the actual data for RL, RS, and $\Delta_4 p$ in the second quarter of 2002 were 0.061, 0.0447 and 0.028, respectively, the corresponding elasticities would be −0.22 (0.061 times −3.65), 0.11 (0.0447 times 2.38), and −0.01 (0.028 times −0.45).

Therefore, *ceteris paribus*, if the short-run interest rate in the second quarter of 2002 increased, say by ten per cent (from 0.0447 to 0.0492), this would have led to a rise of 1.1 per cent in the demand for BM. On the other hand, a similar ten per cent rise in the inflation rate and the rate of return on 10-year Treasury bonds in 2002:2 would have resulted in a 0.1 and 2.2 per cent fall in the demand for BM, respectively. As $|\gamma_2| > |\gamma_3|$, an expected increase in both the cash rate and the rate of return on 10-year Treasury bonds does not have equal effect. If the rate of return on the 10-year Treasury bonds (RL) had increased by x per cent in 2002:2 and the RBA wanted to keep real money balances unchanged, then the cash rate should be raised by two times x per cent because the RL elasticity is twice as larger as the RS elasticity. Consistent with theoretical postulates discussed in Section II, it is also found that an increase in the rate of inflation encourages agents to diversify their portfolios in the economy by acquiring real assets.

Table VIII Empirical results for the short-run demand for BM model $\Delta \ln(m - p - y)_t$

Variable	*Estimated coefficients*	*t-statistics**	*Prob.*	*Expected signs*
Constant	0.002	1.6	[0.11]	+
ΔRL_{t-1}	−0.379	−2.2	[0.03]	−
ΔRS_{t-3}	0.193	2.3	[0.03]	+
ΔRS_{t-4}	0.137	−1.6	[0.11]	+
$\Delta^2 p_t$	−0.409	−4.0	[0.00]	−
$\Delta \ln(m - p - y)_{t-2}$	−0.312	−3.3	[0.00]	+/−
ec_{t-1}	−0.077	−2.5	[0.01]	−
Order of integration of stochastic residuals: I(0)				
$R^2 = 0.33$ when solved for $\Delta \ln(m - p - y)_t$	$F(6,92) = 8$		[0.00]	
$R^2 = 0.962$ when solved for $\ln(m - p - y)_t$				
Diagnostic tests:				
DW	1.86			
AR 1–5:	$F(5,87) = 0.39$	[0.85]		
	$\chi^2(5) = 2.2$	[0.82]		
ARCH 1–4	$F(4,84) = 0.79$	[0.53]		
Normality	$\chi^2(2) = 0.71$	[0.70]		
White heteroskedasticity:				
no-cross terms	$F(12,79) = 0.72$	[0.73]		
cross-terms	$F(27,64) = 0.80$	[0.73]		
RESET	$F(1,91) = 0.03$	[0.85]		

Note: * indicates that the standard errors of coefficients have been corrected by the White Heteroskedasticity-Consistent Standard Errors and Covariance before calculating *t*-ratios.

Using the resulting residuals (the *ec* term) from the long-run relationship in equation (8), one can estimate a VEC model which captures the short-run dynamics of the (inverse) velocity of BM. That is:

$$\Delta(m - p - y)_t = \varphi_0 + \sum_{i=0}^{q_1} \varphi_{1i}\Delta RL_{t-i} + \sum_{i=0}^{q_2} \varphi_{2i}\Delta RS_{t-i} + \sum_{i=0}^{q_3} \varphi_{3i}\Delta\Delta(p)_{t-i} + \sum_{i=1}^{q_4} \varphi_{4i}\Delta(m - p - y)_{t-i} + \theta ec_{t-1} + v_t \quad (9)$$

where φ_{ij} are the estimated short-term coefficients; θ is the feedback effect or the speed of adjustment, whereby short-term dynamics converge to the long-term equilibrium path; and the lagged dependent variables are added to ensure that v_t (or the residual) is white noise. See Hendry, Pagan and Sargan (1984) for a concise discussion of dynamic specification.

Starting with a maximum lag of four for q_1 to q_4, the general-to-specific methodology is now used to omit the insignificant variables in equation (9) on the basis of a battery of maximum likelihood tests. This method of analysis has also been used in other studies. For example see Ericsson, Hendry and Tran (1994), and Hayo (2000). Using I(0) variables in the estimating procedure, joint zero restrictions are imposed on explanatory variables in the general model or equation (9) to obtain the most parsimonious and robust estimators. The empirical results for the parsimonious model capturing short-run dynamics for money demand are presented in Table VIII. As can be seen, the estimated equation for short-run dynamics passes each and every diagnostic test. See Otto (1994) for a concise discussion of diagnostic tests and their importance in the context of the demand for money.

The estimated coefficients have been sensibly signed, with the change in the rate of return on non-financial assets (as proxied by $\Delta^2 p_t$) and the interest rate on assets outside of money (as indicated by the coefficient on ΔRL_{t-1}) having negative elasticities of −0.409 and −0.379, respectively. As expected, changes in the cash rate (ΔRS_{t-3}, and ΔRS_{t-4}) exert a *lagged* positive impact on money demand. Furthermore, the feedback coefficient for the *ec* term is highly significant, validating the significance of the cointegration relationship in the short-run model for money demand. The magnitude of the estimated coefficient for *ec* indicates that the lagged excess money will reduce holdings of money by eight per cent in each quarter.

One problem associated with the analysis of the demand for money is non-constancy of estimated coefficients which can create economic and econometric complications in deriving any inference from the empirical model. Given extensive financial deregulation and innovations introduced in the 1980s, parameter constancy is pivotal in modelling money demand in Australia. Therefore, the estimated short-run model has been evaluated by a number of recursive diagnostic tests which are displayed in Figure 3 in the following order:

$$\begin{bmatrix} a & b & c \\ d & e & f \\ g & h & i \end{bmatrix}$$

where panel (a) displays the recursive residuals; panel (b) depicts the CUSUM test; panel (c) shows the CUSUM of squares; and panels (d) to (i) present the recursively estimated six coefficients (excluding the intercept) over the period 1979:3–2002:2 in the same order that these coefficients appear in Table VIII (from top to bottom). These evaluative tests are useful in assessing the parameter constancy of the model, as recursive algorithms avoid arbitrary splitting of the sample. Overall, the graphical tests reported in Figure 3 reveal that aside from a few minor and insignificant outliers around the 1980s, the test results point to the in-sample constancy of the estimated coefficients. In particular, the recursively estimated coefficients have remained relatively stable since 1985.

IV. Conclusion

After briefly reviewing the relevant literature, this paper determines the long- and short-run drivers of Australia's demand for broad money (BM) using quarterly time series data from 1976:3 to 2002:2. The ADF and KPSS tests for unit roots support the view that all the variables appearing on a standard money demand function are I(1). Therefore, the Johansen cointegration test has been employed to determine the number of the cointegrating vector(s). Cointegration tests clearly indicate that there is a unique cointegrating vector, which links the real demand for BM with real income, the rate of return on 10-year Treasury bonds (*RL*), the official cash rate (*RS*), and the annualised rate of inflation ($\Delta_4 p$).

This paper updates the sample and addresses the problems and shortcomings associated with the previous work on the demand for BM. Unlike previous studies, this paper presents an empirical analysis of the adjustment dynamics of all five variables that enter the cointegrating vector, identifying both the long- and short-term determinants of the demand for BM.

The estimated long-run income elasticity is very close to unity which is consistent with the quantity theory of money and the results obtained in other studies for developed countries, e.g. Beyer (1998) in his study of M3 in Germany, Coenen and Vega (2001) in their recent study of M3 in the Euro area, and Ericsson (1998) in his analysis of the narrow demand for money in the UK. The long-run semi-elasticities of *RL*, *RS* and inflation with respect to the real BM balances are −3.7, 2.4 and −0.45, respectively. The empirical results are broadly in accord with previous studies on the demand for money in developed countries. This paper also presents an error correction model capturing short-run dynamics of money demand. The estimated coefficients

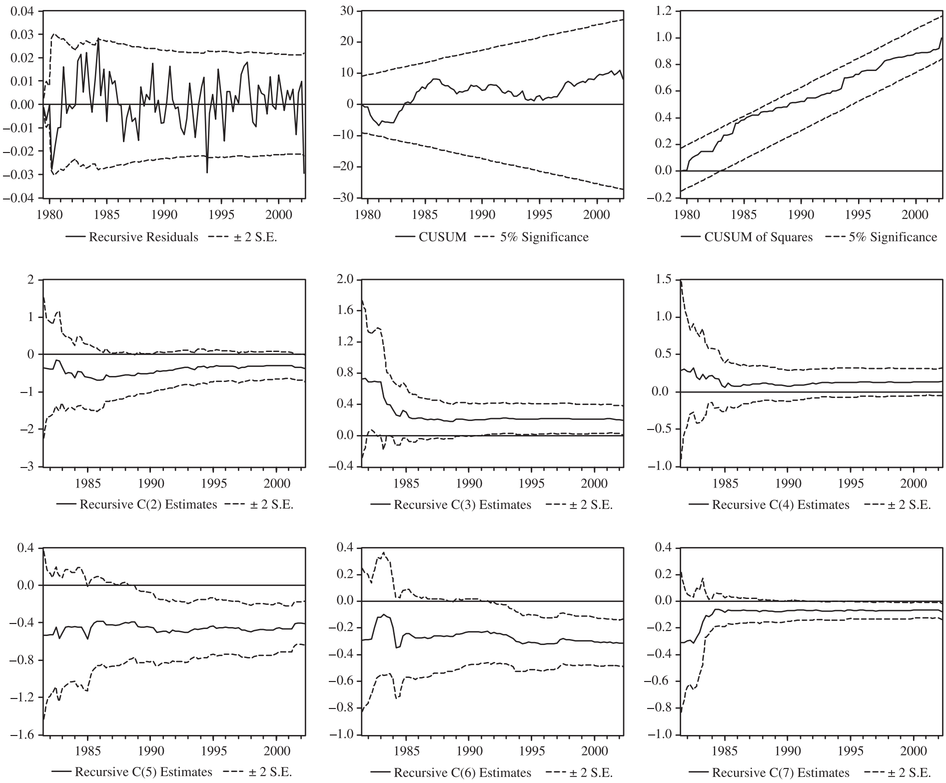

Figure 3. Graphical tests for parameter constancy of the short-run demand for money

Note: One may argue that if stability is equivalent with stationarity in the time series underpinning the broad money target, it will not be necessary to examine ‘regime shifts’ directly. The reported graphical tests above show only in-sample parameter constancy of the estimated coefficients not the absence of any regime shifts.

in this model are not only highly significant but also have consistent signs and orders of magnitude. The estimated error correction model indicates that the selected interest rates adequately represent the prevailing interest rate regime in the economy. This equation shows no sign of misspecification and passes a battery of diagnostic tests.

The major findings of this paper, which can augment our understanding of broad money demand in Australia, are summarised below. First, it is plausible to argue that, *ceteris paribus*, the long- and short-run income elasticities are close to one. Second, inflation has an immediate effect (with no lag) on BM in the short-run, suggesting that an increase in inflation can instantly encourage agents to diversify their portfolios in the economy by acquiring real assets. Third, it seems that a change in the cash rate affects the money demand with three to four quarters lags, whereas the impact of an increase in *RL* on BM is felt after only one quarter. See Table VIII. Therefore, the RBA should pursue a forward looking policy in relation to changes in the cash rate, otherwise the policy may not have a timely and desirable effect. The long- and short-run models estimated for money demand support the view that BM is a predictable monetary aggregate. Fourth, the weak exogeneity test results clearly indicate that unlike y, the three variables RL, RS and Δp are weakly exogenous with respect to $(m - p)$. It is found that the real money gap is an important information variable which has significant explanatory power for output gap and output growth. Given the difficulty of forecasting output gap in Australia this would appear to be a very useful finding.

On this same issue Gerlach and Svensson (2001, p. 24) posit that "it is appropriate to consider both the real money gap and output gap when judging price pressures". In conclusion this paper supports the view that "monetary aggregates" should be used "as a supplementary intermediate target variable in a regime whose principal anchor is an inflation goal" (Laidler, 1999, p. 26).

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