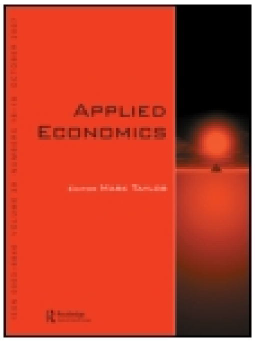

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Australasian money demand stability: application of structural break tests

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Estimates of the demand for money provide important foundations for monetary policy setting but if the estimation technique does not explicitly account for structural changes then such estimates will be biased. This article presents an investigation into the level and stability of money demand (*M1*) for Australia and New Zealand over the period 1960–2009 and demonstrates that both countries experienced regime shifts; Australia also experienced an intercept shift. Application of four time series methods provide consistent results with 1984 and 1998 break dates. Cumulative Sum (CUSUM) and CUSUMSQ stability tests reveal that *M1* demand functions were unstable over the period 1984–1998 for both countries although tests for stability are not rejected thereafter.

Keywords: money demand; cointegration; structural breaks; Australia; New Zealand

JEL Classification: C22; E41

I. Introduction

Empirical analyses of money demand continue with renewed vigour in spite of some established stylized facts concerning income and interest rate elasticities. For advanced countries it is argued that financial reforms introduced in the early 1970s had significant effects on money demand functions and that disequilibrium in money demand functions influenced the effectiveness of interest rate policies in the long run, albeit through its effects on inflation and the output gap. These reforms and the increased use of money substitutes for transactions (e.g. credit/debit cards and electronic money transfers) are argued to have increased competition in financial markets and enhanced international capital mobility. Scale economies in money demand within and across economies may have reduced income elasticities while the contemporaneous utilization of market-based interest rate policies may have improved the rate of interest elasticity.

The choice of monetary policy instrument is crucial; using the incorrect instrument will cause income instability. Deadman and Ghatak (1981) postulated that a stable money demand function is an important issue because it provides a reliable and predictable link between changes in monetary aggregates and changes in variables included in the money

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demand function. Similarly, Poole (1970) argued that the stability aspect of money demand is vital for selecting monetary policy instruments. Explicitly, Poole used Investment Savings-Liquidity Money (ISLM) analysis to show that the money supply (rate of interest) should be targeted if money demand is stable (unstable).

However, even in conditions of stable money demand, many central banks seem to be attracted to targeting the rate of interest following the Taylor rule (see Taylor, 1999). The rationale behind this perspective lies in the belief that adjusting the lagged short term interest rate increases the ability of central banks to influence income and thence central banks now pay less attention to the stability of money demand functions.

Interest rate targeting is a monetary policy framework employed in Australia and New Zealand to stabilize inflation, and such policy selection may be based on either the Taylor rule or a belief that money demand functions are unstable. Although it appears that they have been relatively successful in achieving price stability, their policies have guaranteed neither balanced growth nor macroeconomic stability; this may be due to the added complexities attributable to the liberalization of their financial markets in the 1980s. Financial market liberalization may have caused some instability in the demand for money function which would mean that rate of interest targeting would be the appropriate policy option for central banks. However, reforms and external shocks may have distorted the equilibrium relationship of money demand, and this raises doubts about the validity of studies on money demand that do not utilize structural break estimation methods.

The purpose of this article is to assess the stability of money demand (*M1*) relationship for Australia and New Zealand over the period 1960–2009 while accounting explicitly for structural changes that might have occurred during the period. We apply (i) Lee and Strazicich's (2003) unit root test to test for nonstationarity of the series in the presence of two structural breaks, (ii) Gregory and Hansen's (GH, 1996a, b) single endogenous break test to test for cointegration among the variables and to estimate the cointegrating equations. Standard time series techniques of (iii) Hendry's General to Specific (GETS), (iv) Engle and Granger's (1987) two step method (EG), (v) Phillips and Hansen's (1990) Fully Modified Ordinary Least Squares (FMOLS) and (vi) Two Stage Least Squares (2SLS) are then applied to conduct sub-sample period estimations. This article has the following structure. Section II presents a review of the literature. The methods and empirical results are detailed in Sections III and IV, respectively. Conclusions are provided in Section V.

II. Brief Review of Time Series Studies

Although there is a vast literature that presents investigations into the level and stability of money demand using cross-section, time-series or panel data estimation methods, many of the results are neither totally consistent across studies nor based on estimation methods that explicitly allow for structural breaks in the time series relationships. This is exemplified by recent studies on money demand that relate to advanced countries and, more specifically, to Australia and New Zealand.[1]

Advanced countries

The stability of money demand functions has been widely researched. Hoffman *et al*. (1995) constrained the income elasticity to be unity when analysing post-war data (1955–1990) and provided evidence which suggests that *M1* demand is stable in Canada, Japan, the UK, the USA and West Germany. Lutkepohl and Wolters (1998) analysed the *M3* demand relationship for Germany over the period 1976–1996 and corroborates stability when the income elasticity was constraint at unity. Similar results were obtained by Maki and Kitasaka (2006) and Lucas (1988) for Japan and USA, respectively.

Studies that estimated unconstrained income elasticities include Artis *et al*. (1993) who identified significant income elasticities around 1.2 for *M1* and *M2* demand for Belgium, Denmark, France, Germany, Ireland, Italy and the Netherlands between 1979 and 1990; similar estimates were attained by Monticelli and Strauss-Kahn (1993). The often found income elasticity above unity is explained within the standard portfolio approach by the neglect of a wealth variable in the cointegrating vector. When Ewing and Payne (1999) examined *M1* demand for Australia, Austria, Canada, Finland, Italy, Germany, Switzerland, the UK and the USA they identified a range of income elasticities between 0.5 and 1.2 and suggest that *M1* demand was stable in Australia, Austria, Finland, Italy, the UK and the USA when *M1* is cointegrated with real income and the nominal interest rate; stability was identified for Canada,

[1] For discussions related to the theoretical developments of the demand for money see Laidler (1969, 1977, 1993a, b), Barnett *et al*. (1992), Bruggemann and Nautz (1997), Serletis (2001) and Duca and van Hoose (2004).

Germany and Switzerland also but only when the exchange rate was incorporated. Baba *et al.* (1992) estimated the demand for *M1* for USA over the period 1960–1988 and obtained an income elasticity of around 0.5; comparable results for USA were obtained by Ball (2001) and Choi and Jung (2009). Clearly there is dispute over the income elasticity estimate as Haug and Lucas (1996) also examined *M1* demand for Canada over the period 1953–1990 and attained an income elasticity of around 0.4, while similar findings for Canada were obtained by Georgopoulos (2000).[2] In spite of the large variation in income elasticity estimates the aforementioned studies either implicitly or explicitly support central banks' monetary targeting regimes.

However, efforts by Bahmani-Oskooee and Chomsisengphet (2002) suggest that money demand is not universally stable. They assessed the stability of *M2* demand for 11 Organization for Economic Co-operation and Development (OECD) countries and obtained a range of income elasticities between 0.6 and an implausibly high 3.9. Although their findings indicate that money demand is stable in Australia, Austria, Canada, France, Italy, Japan, Norway, Sweden and the USA, and they also suggest some *instability* of *M2* for Switzerland and the UK. Obtaining evidence against the stability of money demand suggests that interest rate targeting is optimal.

Corroborating evidence for money demand instability is not unheard of. For Canada, both McPhail (1991) and Haug (1999) asserted that the openness of financial systems had made significant impacts on broader monetary aggregates and therefore support interest rate targeting. Similarly, Nagayasu (2003) obtained a near-unit income elasticity estimate of *M2* demand for Japan over the period 1958–2000 and, through application of Hansen's (1992) stability tests, revealed that *M2* demand is unstable.

Papadopoulos and Zis (1997) investigated the determinants and the stability of money demand (*M1*, *M2* and *M3*) for Greece. Although they find that *M2* and *M3* are largely stable, they also obtain results which suggest that *M1* demand is unstable; this corroborates earlier findings of Sharma (1994). In a study of the Spanish economy, Vega (1998) finds that a structural break, which may capture changes in the openness of the financial system, has affected the stability of broad money. This leads Vega (1998) to suggest that it is reasonable to use the rate of interest to curtail inflation rates.[3] Other recent studies, such as Coenen and Vega (2001), Bruggeman *et al.* (2003), Brand and Cassola (2004), Beyer (2009) and Belke and Czudaj (2010) all provide useful inferences on the Euro Area money demand and monetary policy.

The case of Australia

The pioneering study by Cohen and Norton (1969) implied stability in narrow and broad measures of money. Their study was replicated and augmented by others for various monetary aggregates. Corroborating evidence was provided by Sharpe and Volker (1977) and Pagan and Volker (1981) who found limited instability in money demand functions. Hoque and Al-Mutairi (1996) investigated the long-run relationship between *M1* and its determinants (income, interest rate and price level) over the period 1970–1993 and found no instability in *M1* demand despite the countenance of financial innovation and deregulation. Valadkhani (2005) examined the determinants of *M2* demand over the period 1976–2002 and found it to be cointegrated with real income, the rate of return on 10-year Treasury bonds, and cash and inflation rates, with an income elasticity of *M2* demand close to unity. Felmingham and Zhang (2001) examined *M2* demand over the period 1976–1998 and found it to be stable subject to a regime shift occurring during the 1991 recession, which supported earlier findings by Lim and Martin (1991), Juselius and Hargreaves (1992), Lim (1995) and Asano (1999). However, Felmingham and Zhang (2001) attained an implausibly high income elasticity of 1.2; much lower income elasticity is expected due to increased financial efficiencies and scale economies in money demand.

Sets of empirical results that question the stability of money demand in Australia include Felmingham and Zhang (2001), who found some instability in the 1990s, and Adams and Porter (1976) and Blundell-Wignall and Thorp (1987) who both provided evidence that led them to argue against the stability of narrow and broad monetary aggregates. Orden and Fisher (1993) examined the dynamic impacts of financial deregulation on *M3* demand over the period 1965–1989 and found a cointegrating relationship between real *M3* and prices and output series prior to the financial liberalization; however they did not support cointegration between *M3* demand and its

[2] Other studies that found no evidence of instability in money demand functions include Hayo (2000) for Austria, Juselius (1998) for Denmark, Nielsen *et al.* (2004) for Italy, Bahmani-Oskooee and Economidou (2005) for Greece, Gerlach-Kristen (2001) for Switzerland, and Nielsen (2004) and Escribano (2004) for the UK.

[3] On a policy front, Papadopoulos and Zis (1997) are doubtful whether a monetary rule can provide an efficient anti-inflation policy framework.

price and output determinants either over the full sample or after financial liberalization, and this implies instability in the *M3* demand function over the entire period and especially subsequent to 1982.

The case of New Zealand

There is a dearth of empirical studies on money demand for New Zealand and the stability of her various monetary aggregates is yet to be determined. Orden and Fisher (1993) found some instability of money demand in Australia; however their results for New Zealand are different as they found stability over the whole and sub-periods. Siklos (1995a, b) examined the cointegrating links between *M3*, expected inflation and short term interest rates (the difference between *NZ* and *US* rates) over the period 1981–1994 and attained implausibly high income elasticities varying between 2 and 6. The income elasticities attained by Choi and Oxley (2004) and Valadkhani (2002) also seem unexpectedly high at around 1.7 and 1.5, respectively. An income elasticity estimate that is more in line with expectations was provided by Razzak (2001) who found that the income elasticity of monetary base to be around unity over the period 1988–1997 while asserting that the correlation between money and real output is stronger than that between money and inflation.

Empirical issues

Given that a number of major financial reforms were implemented by Australia and New Zealand since the 1960s to enhance the efficiency of their financial sectors, it is entirely plausible that structural changes in their money demand may have occurred. Moreover, other events that influenced their domestic economies (such as natural disasters, oil price shocks and financial crises, etc.) may be associated with structural changes in the data series also. The failure to accommodate structural changes in the data series and cointegrating vectors could result in the attainment of misleading results.

Although the aforementioned Australia and New Zealand studies offer important insight on monetary policy procedures, their empirical results are neither mutually supportive nor equivocal. Furthermore, with the notable exception of Felmingham and Zhang (2001) for Australia (albeit with an implausibly high income elasticity), most studies used standard time series methods that allow for no formal tests of structural breaks.

From the early 1980s, both countries underwent continuing economic liberalisation. In Australia, the mid-1980s saw financial deregulation and the Australian dollar float, while in 2000 the introduction of a Goods and Services Tax (GST) sought to encourage savings amongst low income earners. The formation of the Australian Stock Exchange Limited in 1987 and microeconomic reforms in the manufacturing sector both boosted private investment. Similarly, a number of events also affected New Zealand's economic performance; for instance, she lost her preferential trading position with the UK in 1973, embarked on financial market deregulations in the 1980s, undertook privatization measures during 1980s and 1990s, and experienced the Asian financial crises and climate drought in the late 1990s.

The article fills this gap in the literature by presenting estimates of the demand for money (*M1*) for Australia and New Zealand over the period 1960–2009. Structural breaks in the data series and cointegrating vectors are examined through the use of Lee and Strazicich (2003, 2004) and GH (1996a, b) methods; naturally Felmingham and Zhang (2001) were only able to apply the latter of these two methods.

III. Specification and Methods

Conventionally the demand for money is specified as a function of real income and the nominal interest rate, however to capture the true cost of holding money we specify the demand for money in its canonical form and its extended versions, such that

$$\ln M_t = \theta_0 + \theta_y \ln Y_t + \theta_R R_t + \varepsilon_t \tag{1}$$

$$\ln M_t = \theta_0 + \theta_y \ln Y_t + \theta_R R_t + \theta_E \ln E_t + \theta_\pi \pi_t + \varepsilon_t \tag{2}$$

where θ_0 = intercept, M = real narrow money stock, Y = real output, R = cost of holding money proxied with the nominal short-term interest rate, E = cost of holding money proxied with the real effective exchange rate, π = cost of holding money proxied with the inflation rate and $\varepsilon \approx N(0, \sigma)$. Real money balances are defined as the narrow monetary aggregate, *M1*, deflated by the Gross Domestic Product (GDP) deflator.[4] Real output is constructed using

[4] Many central banks, including the Reserve Bank of Australia and Reserve Bank of New Zealand, find it relatively easy to control *M1* and therefore testing for the stability of *M1* demand offers useful implications on monetary policy procedures. Although either nominal or real exchange rate can be used to proxy for the cost of holding money, we have used real effective exchange rate due to data availability. Our results are based on the application of the GDP deflator to compute the inflation rate although application of the Consumer Price Index (CPI) gave qualitatively similar results.

nominal GDP (deflated by the GDP deflator) and the change in the GDP deflator is our proxy of the inflation rate. The 3-month deposit rate is our proxy for the nominal interest rate. Annual data for the period 1960–2009 were obtained from International Financial Statistics (2010) and the World Development Indicators (World Bank, 2010).[5]

Our explicit expectations of the sign and magnitude of the real income variable is in line with Baumol-Tobin and Quantity Theory models which predict that the income elasticity should be 0.5 and 1, respectively (Baumol, 1952; Friedman, 1956; Tobin, 1956). Ball (2001) pointed out that low income elasticity estimates would imply that the Friedman rule is not optimal and that the money supply should grow more sluggishly than income to attain price stability. In advanced countries, the income elasticity is expected to be much lower than unity due to improvements in and developments of financial systems. Our explicit expectations of the signs and magnitudes of cost of holding money variables (nominal interest rate, inflation rate and real effective exchange rate) are negative and small.[6]

Lee and Strazicich (2003) tests

The endogenous two-break Lagrange Multiplier (LM) unit root tests proposed by Lee and Strazicich (2003) can be explained using two models, e.g. model A and model C. Both models are based on alternative assumptions about structural breaks; model A allows for two shifts in the intercept and model C includes two shifts in the intercept and trend.

Model A is specified as follows:

$$Z_t = [1, t, D_{1t}, D_{2t}]' \tag{3}$$

where $D_{jt} = 1$ for $t \geq T_{Bj} + 1$, $j = 1, 2$, and 0 otherwise. The break date is denoted by T_{Bj}. The null and alternative hypotheses of model A are

$$\begin{aligned} H_0\colon\ y_t &= \mu_0 + d_1 B_{1t} + d_2 B_{2t} + y_{t-1} + \nu_{1t} \\ H_1\colon\ y_t &= \mu_1 + \gamma t + d_1 D_{1t} + d_2 D_{2t} + \nu_{2t} \end{aligned} \tag{4}$$

The specification, null and alternative hypotheses of model C, respectively, are

$$Z_t = [1, t, D_{1t}, D_{2t}, DT_{1t}, DT_{2t}]' \tag{5}$$

$$\begin{aligned} H_0\colon\ y_t &= \mu_0 + d_1 B_{1t} + d_2 B_{2t} + d_3 D_{1t} + d_4 D_{2t} + y_{t-1} + \nu_{1t} \\ H_1\colon\ y_t &= \mu_1 + \gamma t + d_1 D_{1t} + d_2 D_{2t} + d_3 DT_{1t} + d_4 DT_{2t} + \nu_{2t} \end{aligned} \tag{6}$$

where $DT_{jt} = t - T_{Bj}$ for $t \geq T_{Bj} + 1$, $j = 1, 2$, and 0 oterwise; $B_{jt} = 1$ for $t = T_{Bj} + 1$, $j = 1, 2$, and 0 otherwise; ν_{1t} and ν_{2t} denote the stationary error terms. The LM unit root test statistic can be obtained by estimating

$$\Delta y_t = \delta' \Delta Z_t + \phi \overline{S}_{t-1} + \mu_t \tag{7}$$

where $\overline{S}_t = y_t - \overline{\psi}_x - Z_t \overline{\delta}$, $t = 2, \ldots, T$; Δy_t is just regressed on ΔZ_t to provide estimates of $\overline{\delta}$; $\overline{\psi}_x = y_1 - Z_1 \overline{\delta}$ and the first observations of y_t and Z_t are y_1 and Z_1, respectively. The LM test statistics are provided by $\overline{\tau}$ which is the test statistic for the unit root null hypothesis that $\phi = 0$.

Initially we allocated a maximum lag length of eight periods and obtained the optimal lag length on the basis of the significance of the last lag. The break dates are determined where the LM test statistic is at its minimum. The critical values of this test are tabulated in Lee and Strazicich (2003, 2004). Thus this method is more demanding than Perron (1989, 1997) because it offers more than one break in the series.

Gregory and Hansen tests

Unlike the Bai and Perron (2003) and Lee and Strazicich (2003) tests, GH's (1996a, b) method is a test for structural changes in the cointegrating vector. The null hypothesis of no cointegration with structural breaks is tested against the alternative of cointegration. GH has postulated four models that are based on alternative assumptions about structural breaks: *model 1* is a level shift; *model 2* is a level shift with trend; *model 3* is a regime shift where both the intercept and the slope coefficients change and *model 4* is a regime shift where intercept, trend and slope coefficients all change. The single break date in these models is endogenously determined. Based on Equation 2, the implied specification of these four models with structural breaks, respectively, are as follows:

$$\begin{aligned} \ln M_t = \mu_1 &+ \mu_2 \varphi_{tk} + \alpha_1 \ln Y_t + \alpha_2 R_t \\ &+ \alpha_3 \ln E_t + \alpha_4 \pi_t + \varepsilon_t \end{aligned} \tag{8}$$

$$\begin{aligned} \ln M_t = \mu_1 &+ \mu_2 \varphi_{tk} + \beta_1 t + \alpha_1 \ln Y_t \\ &+ \alpha_2 R_t + \alpha_3 \ln E_t + \alpha_4 \pi_t + \varepsilon_t \end{aligned} \tag{9}$$

[5] Using annual data our dataset is balanced and consequently there are no gaps in the series. Selecting monthly or quarterly data would have resulted in attaining data from 1990 onwards and it would have been difficult to analyse the impact of reform policies implemented during 1980s.

[6] See Laidler (1993a, b), Sriram (1999) and Hoffman and Rasche (2001) for surveys of long run elasticities of money demand.

Table 1. Two-break minimum LM unit root test, 1960–2009

	Australia				New Zealand			
	Model A		Model C		Model A		Model C	
Variables	Test statistic	Break dates	Test statistic	Break dates	Test statistic	Break dates	Test statistic	Break dates
$\ln M$	−2.003 [4]	1981; 2005	−1.047 [2]	1984; 1986	−0.263 [3]	1987; 1998	−3.987 [3]	1984; 1986
$\ln Y$	−3.112 [5]	2000; 2003	−0.182 [6]	2003; 2004	−1.237 [4]	1973; 1984	−2.376 [3]	1995; 2003
R	−1.280 [7]	1984; 1988	−1.601 [5]	1981; 1985	−2.128 [5]	1975; 1986	−3.228 [6]	1997; 2002
$\ln E$	−2.152 [6]	1987; 1995	−2.251 [4]	1987; 1988	−2.187 [4]	1986; 1992	−3.721 [5]	1988; 1991
π	−2.120 [5]	2000; 2002	−2.672 [4]	2001; 2002	−3.036 [2]	1984; 1987	−1.121 [4]	1995; 2003

Notes: The 5% critical values for Models A and C are −3.842 and −5.286, respectively. The number in square brackets indicates the optimal number of lagged first-differenced terms included in the unit root test to correct for serial correlation. Critical values are taken from Lee and Strazicich (2003, 2004).

$$\ln M_t = \mu_1 + \mu_2\varphi_{tk} + \beta_1 t + \alpha_1 \ln Y_t + \alpha_{11} \ln Y_t\varphi_{tk} + \alpha_2 R_t + \alpha_{22} R_t\varphi_{tk} + \alpha_3 \ln E_t + \alpha_{33} \ln E_t\varphi_{tk} + \alpha_4\pi_t + \alpha_{44}\pi_t\varphi_{tk} + \varepsilon_t \quad (10)$$

$$\ln M_t = \mu_1 + \mu_2\varphi_{tk} + \beta_1 t + \beta_2 t\varphi_{tk} + \alpha_1 \ln Y_t + \alpha_{11} \ln Y_t\varphi_{tk} + \alpha_2 R_t + \alpha_{22} R_t\varphi_{tk} + \alpha_3 \ln E_t + \alpha_{33} \ln E_t\varphi_{tk} + \alpha_4\pi_t + \alpha_{44}\pi_t\varphi_{tk} + \varepsilon_t \quad (11)$$

where μ_1 is the intercept, μ_2 is the parameter for intercept shift, α_1 is the income elasticity, β_1 is the parameter for trend, β_2 is the parameter for trend shift, α_2 is the semi-rate of interest elasticity, α_3 is the exchange rate elasticity, α_4 is the elasticity with respect to inflation rate, α_{11} is the parameter for shift in income elasticity, α_{22} is the parameter for shift in semi-rate of interest elasticity, α_{33} is the parameter for shift in exchange rate elasticity and α_{44} is the parameter for shift in elasticity with respect to the inflation rate. A break date is selected where the absolute value of the Augmented Dickey–Fuller (ADF) test statistic is at its maximum. The critical values for cointegration are tabulated in GH (1996a, b) and are used for testing cointegration in the *EG* method with unknown breaks.[7]

IV. Empirical Results

Lee and Strazicich (2003) tests

Endogenous two break minimum LM unit root tests were applied to assess the order of integration of variables. Table 1 reports the results for these tests based on models A and C which represent two breaks in the intercept (model A) and two breaks in the intercept and trend (model C). The test statistics of the LM unit root tests for the five variables (real *M1*, real income, nominal interest rate, real effective exchange rate and inflation rate) do not exceed the critical values in absolute terms and therefore the unit root null cannot be rejected at the 5% level. The *t*-statistics corresponding to the break dates are statistically significant at conventional levels (not reported for brevity). Break dates are fairly consistent across models, are expected for both countries and are in line with the timings of macroeconomic events outlined above.

Cointegration tests

The GH method was applied to test for cointegration between the variables in canonical and extended equations of money demand (i.e. Equations 1 and 2, respectively); results are provided in Table 2. The null hypothesis of no cointegration is rejected for canonical specification (1) in models 1 (break date (hereafter BD): 1994) and 4 (BD: 1984) for Australia and in models 3 (BD: 1998) and 4 (BD: 1984) for New Zealand. For specification (2), models 1 and 2 reject the null hypothesis of no cointegration for Australia and the break dates are 1984 and 1997, respectively. Using the same specification, the null hypothesis of no cointegration is rejected only in model 4 for New Zealand with a break date of 1984. These results support the existence of long-run relationships of the demand for money in both countries. Explicitly, the results of the canonical form show that money demand is cointegrated with real income and the nominal interest rate; the same can be observed when the model is augmented with real effective exchange and inflation rates, as in the extended version.

[7] GH developed the critical values by modifying the MacKinnon (1991) procedure.

Table 2. Cointegration tests with structural breaks, 1960–2009

	Specification/GH model	Break date	GH test statistic	5% critical value	Existence of cointegration
Australia	*Canonical specification*				
	Model 1	1994	−5.036	−3.190	Yes
	Model 2	2000	−1.754	−3.190	No
	Model 3	1987	−0.306	−3.190	No
	Model 4	1984	−4.667	−3.190	Yes
	Extended specification				
	Model 1	1984	−3.972	−3.603	Yes
	Model 2	1997	−9.116	−3.603	Yes
	Model 3	1982	−1.734	−3.190	No
	Model 4	2001	−2.062	−3.190	No
New Zealand	*Canonical specification*				
	Model 1	1984	−1.673	−3.603	No
	Model 2	1986	−2.996	−3.603	No
	Model 3	1998	−6.387	−3.603	Yes
	Model 4	1984	−7.900	−3.603	Yes
	Extended specification				
	Model 1	1988	−2.370	−3.190	No
	Model 2	1984	−2.776	−3.190	No
	Model 3	2005	−0.062	−3.190	No
	Model 4	1984	−8.024	−3.603	Yes

Break dates for both countries are consistent with those attained through the application of Lee and Strazicich's (2003) method. A majority of the break dates are in 1980s; this is not unexpected because both countries underwent major economic reforms in the 1980s and the break dates may highlight the importance of financial reforms in these domestic economies.

We apply the GH method based on the premise that structural breaks may have affected the cointegrating relationships of money demand in both countries. Strictly speaking, structural break tests should only be used when standard methods fail to yield robust estimates. Applications of the standard Johansen (1991, 1995) method to data for the whole time period did not give meaningful results; see Table A1 and A2 in the Appendix. Note that the results obtained from application of the Johansen method did reveal weak cointegration among the variables in canonical and extended equations for Australia and unexpectedly high income elasticity estimates of around 1.8 (canonical equation) and 2.1 (extended equation) for New Zealand. In light of these Johansen results, we argue that there could be structural breaks present in the *M1* relationships for both countries and therefore our application of the GH method is justifiable.

Long run estimates

GH cointegrating equations were estimated with the EG method and the results are presented in Table 3. Given *a priori* expectation that the income elasticity estimates should be less than unity, we can conclude that there are plausible results for Australia in model 4 (canonical specification) and model 1 (extended specification) and plausible results for New Zealand in model 4 (extended specification). The estimated coefficients in these models have expected signs and are statistically significant at the 95% confidence level. For Australia, the income elasticity of money demand is around 0.64, which implies that a 1% increase in real income raises the demand for money by about 0.64%, while for New Zealand the income elasticity of money demand is around 0.68, which implies that a 1% increase in real income would raise the demand for money by about 0.68%, all *ceteris paribus*.[8] With these findings, we argue that the money demand relationships in Australia and New Zealand have undergone regime shifts where intercept, trend and slope coefficients have changed. Australian money demand has also undergone both intercept shift (extended specification) and regime shift (canonical specification) with the latter appearing to be dominant.

[8] We disregarded the estimates of other models for both countries because they are either statistically insignificant or have implausible income elasticity magnitudes. The canonical specification failed to explain the determinants of money demand for New Zealand, leading us to prefer the extended version.

Table 3. GH cointegrating equations, 1960–2009

	Canonical specification				Extended specification		
	Model 1 (Aus)	Model 4 (Aus)	Model 3 (NZ)	Model 4 (NZ)	Model 1 (Aus)	Model 2 (Aus)	Model 4 (NZ)
C	1.367	0.662	2.370	6.977	0.890	−3.467	1.028
	(2.18)*	(2.26)*	(0.76)	(3.26)*	(3.26)*	(0.77)	(6.87)*
Dum × C	−0.322	−1.263	1.277	−0.283	−0.214	−1.273	−0.552
	(1.26)	(2.55)*	(0.28)	(1.24)	(5.62)*	(0.54)	(1.96)*
T	–	0.002	0.170	0.161	–	0.334	0.898
		(7.85)*	(1.50)	(2.34)*		(2.31)*	(4.87)*
Dum × T	–	−0.273	–	0.332	–	–	1.256
		(3.41)*		(1.50)			(5.05)*
ln Y_t	2.560	0.643	−1.231	3.277	0.635	5.661	0.679
	(0.25)	(4.76)*	(1.18)	(1.61)	(4.29)*	(1.03)	(3.12)*
Dum × ln Y_t	–	0.541	−0.788	3.421	–	–	0.530
		(2.07)*	(0.86)	(0.69)			(4.00)*
R_t	−0.162	−0.047	−1.259	−0.887	−0.067	−0.135	−0.015
	(1.24)	(5.23)*	(1.26)	(1.52)	(2.60)*	(4.23)*	(2.46)*
Dum × R_t	–	−0.011	−0.323	−0.162	–	–	−0.008
		(1.99)*	(0.13)	(0.89)			(2.01)*
ln E_t	–	–	–	–	−0.099	−0.350	−0.104
					(5.64)*	(0.76)	(4.37)*
Dum × ln E_t	–	–	–	–	–	–	−0.087
							(1.75)**
π_t	–	–	–	–	−0.102	−2.345	−0.045
					(3.01)*	(1.22)	(3.03)*
Dum × π_t	–	–	–	–	–	–	−0.020
							(1.80)**

Notes: Aus and NZ means Australia and New Zealand, respectively. Absolute *t*-ratios are in parentheses. C and T denote intercept and trend, respectively. Dummy variables are created using the break dates; for example, in canonical specification model 1 for Australia the break date is 1994 therefore dummy is unity after 1994.
* and ** indicate significance at 5 and 10% levels, respectively.

Sub-sample estimates

Given the presence of these obtained break dates it is prudent to examine long run elasticities of money demand for sub-sample periods.[9] The observed common break is 1984, and moreover a break in late 1990s is also present for both countries. Consequently we select two sets of sub-samples such that pre-reforms periods are 1960–1983 and 1960–1997 and post-reform periods are 1984–2009 and 1998–2009. The break date in 1984 is not unrealistic because both countries implemented financial reforms around that time. Further, the break date in 1998 could also be justified as Australia launched unilateral trade liberalization measures and internal structural reforms during the 1990s which led to higher rates of growth of GDP and productivity. Some examples of these reforms include tariff reform and deregulation and privatization of many services sectors. Similarly the New Zealand economy was also affected by a number of economic events that took place during late 1990s, such as the 1996 and 1998 income tax reforms in, the 1997 Asian financial crises and several state enterprise privatizations.

Application of four time series methods viz., GETS, EG, FMOLS and 2SLS give consistent results for both sets of sub-samples;[10] see Table 4 and 5 for the sub-sample cointegrating equations based on canonical and extended equations, respectively. The estimated coefficients have expected signs and are significant at conventional levels. Almost without exception, the income elasticity estimates are less than unity and the estimates of nominal interest, real effective exchange and inflation rates have the expected negative signs. Following Engle and Granger (1987) we also tested for the stationarity of the resulting EG residuals for the sub-sample periods. Applications of the ADF unit root test show that in all cases the residuals are stationary, thereby corroborating the cointegration case. (The ADF unit root test results for the residuals are reported in Table A3 in the Appendix.) A smaller number of observations

[9] We only considered break dates of those models which reveal the existence of cointegration.
[10] See Kumar *et al*. (2012a, b) and Rao (2007) for details on alternative time series methods.

Table 4. Cointegrating equations for sub-sample periods; canonical specification

	ln Y (Aus)	R (Aus)	ln Y (NZ)	R (NZ)
GETS				
1960–1983	0.867	−0.086	0.890	−0.023
	(2.33)*	(1.97)*	(7.54)*	(2.45)*
1984–2009	0.651	−0.103	0.717	−0.765
	(3.20)*	(4.35)*	(3.47)*	(1.87)**
1960–1997	0.803	−0.072	0.856	−0.176
	(4.45)*	(2.58)*	(4.35)*	(2.36)*
1998–2009	0.752	−0.099	0.785	−0.189
	(2.12)*	(2.00)*	(5.32)*	(2.89)*
EG				
1960–1983	0.892	−0.120	0.853	−0.009
	(2.54)*	(2.06)*	(3.24)*	(2.60)*
1984–2009	0.670	−0.167	0.652	−0.102
	(2.30)*	(3.25)*	(3.07)*	(2.01)*
1960–1997	0.866	−0.009	0.843	−0.086
	(6.73)*	(1.68)**	(3.85)*	(2.33)*
1998–2009	0.710	−0.024	0.802	−0.105
	(4.50)*	(1.85)**	(4.01)*	(1.70)**
FMOLS				
1960–1983	0.923	−0.092	0.844	−0.068
	(2.87)*	(1.69)**	(3.70)*	(2.39)*
1984−2009	0.697	−0.103	0.723	−0.239
	(2.34)*	(3.95)*	(3.56)*	(1.71)**
1960–1997	0.899	−0.016	0.801	−0.122
	(3.04)*	(1.76)**	(1.89)**	(2.04)*
1998–2009	0.778	−0.018	0.795	−0.126
	(3.20)*	(1.68)**	(1.75)**	(1.98)*
2SLS				
1960–1983	0.958	−0.177	1.026	−0.340
	(1.90)**	(2.69)*	(1.79)**	(2.42)*
1984−2009	0.693	−0.181	0.802	−0.389
	(2.56)*	(1.80)**	(2.05)*	(1.78)**
1960–1997	0.870	−0.024	0.962	−0.095
	(2.37)*	(2.16)*	(1.67)**	(2.23)*
1998–2009	0.791	−0.029	0.831	−0.101
	(4.04)*	(1.82)**	(1.69)**	(2.37)*

Notes: Absolute t-ratios are in parentheses. Aus and NZ signifies Australia and New Zealand, respectively.
* and ** indicate significance at 5 and 10% levels, respectively.

for the sub-samples raise the concern for endogeneity and short sample biases, however, according to Rao (2007) if alternative time series methods give consistent cointegrating estimates then the aforementioned issues are minimal.

The sub-sample estimates provide useful insight on whether the financial reforms had any significant effect. If they have been effective then there should be evidence for some economies of scale in the use of *M1*; further the response of the demand for money to the rate of interest should improve because of a progression towards more market-based interest rate policies and increased capital mobility. In other words and relative to the pre-reform period, the post-reform sub-samples should show a relatively lower income elasticity estimate while the absolute value of the interest rate estimate should increase.

The results in Table 4 and 5 show that income (interest rate) elasticities in both canonical and extended equations have declined (increased) in the post-reform sub-samples. Further, in most cases the estimates of real effective exchange and inflation rates have increased relative to the pre-reform estimates. These observed changes in the long run elasticities seem to be slightly greater in the first set of sub-samples where the break date is 1984, and they may be illustrating that reforms have improved the financial efficiency in both countries. Also, it is likely that structural breaks may have caused some short-run instability in the money demand functions.

Short run estimates

The short run error correction models (ECM) are estimated with Hendry's GETS approach[11] with the GH cointegrating equations used to establish the ECM models. The dependent variable ($\Delta \ln M_t$) is regressed on its lagged values, the current and lagged values of explanatory variables ($\Delta \ln Y_t$, ΔR_t, $\Delta \ln E_t$ and $\Delta \pi_t$) and the one period lagged residuals from the respective GH cointegrating equation. Application with a maximum of four period lags and further application of variable deletion tests provide parsimonious ECM models, as reported in Table 6. Two ECM models are estimated using Australian data, based on GH models 1 and 4 and presented in columns Aus (1) and Aus (2); the results of the ECM model based on New Zealand data, which are based on GH model 4, are presented in column NZ (1).

The short run dynamic estimates are statistically significant at the 5% level and the lagged error correction term (ECM_{t-1}) has the expected negative sign; this implies a negative feedback mechanism which suggests that if there are departures from equilibrium in the previous period then this departure is reduced in the current period by about 21%–25% for Australia and by about 11% for New Zealand.[12]

[11] See Taylor (1986) and Rao *et al.* (2010) for an overview and strengths of the GETS technique.
[12] The χ^2 statistics indicate that there are no diagnostic test issues associated with serial correlation (χ^2_{sc}), functional form misspecification (χ^2_{ff}), nonnormality (χ^2_n) or heteroscedasticity (χ^2_{hs}) in the residuals; hence, the short run dynamic results are well-determined and robust.

Table 5. Cointegrating equations for sub-sample periods; extended specification

	ln Y (Aus)	R (Aus)	ln E (Aus)	π (Aus)	ln Y (NZ)	R (NZ)	ln E (NZ)	π (NZ)
GETS								
1960–1983	0.876	−0.180	−0.265	−0.071	0.885	−0.005	−0.820	−0.553
	(2.74)*	(1.64)**	(1.68)**	(2.34)*	(4.35)*	(2.67)*	(5.46)*	(1.67)**
1984–2009	0.664	−0.231	−0.179	−0.102	0.703	−0.103	−1.067	−0.871
	(2.79)*	(2.05)*	(2.24)*	(1.87)**	(3.74)*	(1.99)*	(3.28)*	(1.70)**
1960–1997	0.889	−0.085	−1.087	−0.421	0.900	−0.096	−0.127	−0.162
	(2.36)*	(2.40)*	(1.70)**	(3.45)*	(1.76)**	(2.74)*	(1.80)**	(2.51)*
1998–2009	0.732	−0.103	−0.121	−0.113	0.843	−0.105	−0.134	−0.239
	(2.60)*	(2.59)*	(2.05)*	(1.66)**	(2.04)*	(2.29)*	(1.89)**	(1.87)**
EG								
1960–1983	0.873	−0.076	−0.016	−0.082	0.972	−0.026	−0.273	−0.120
	(2.32)*	(3.25)*	(1.80)**	(2.60)*	(2.35)*	(2.30)*	(2.76)*	(4.25)*
1984–2009	0.612	−0.189	−0.210	−0.112	0.655	−0.135	−0.821	−0.237
	(2.05)*	(2.43)*	(2.07)*	(1.64)**	(3.91)*	(3.29)*	(1.88)**	(1.69)**
1960–1997	0.874	−0.021	−0.127	−0.062	0.921	−0.011	−0.062	−0.028
	(2.88)*	(2.37)*	(2.16)*	(1.70)**	(4.36)*	(2.82)*	(2.36)*	(2.73)*
1998–2009	0.718	−0.175	−0.188	−0.100	0.835	−0.082	−0.283	−0.184
	(1.98)*	(2.31)*	(1.66)**	(1.90)**	(2.52)*	(1.75)**	(3.03)*	(1.79)**
FMOLS								
1960–1983	1.073	−0.008	−0.190	−0.002	0.890	−0.133	−0.006	−0.122
	(2.67)*	(3.87)*	(2.60)*	(1.67)**	(1.85)**	(2.08)*	(2.06)*	(4.25)*
1984–2009	0.734	−0.056	−0.197	−0.025	0.751	−0.205	−0.133	−0.207
	(2.29)*	(1.65)**	(2.92)*	(1.84)**	(2.11)*	(1.79)**	(1.69)**	(2.21)*
1960–1997	0.974	−0.040	−0.134	−0.012	0.673	−0.016	−0.026	−0.016
	(2.27)*	(3.23)*	(2.74)*	(2.37)*	(3.37)*	(2.14)*	(2.93)*	(2.91)*
1998–2009	0.705	−0.104	−0.189	−0.333	0.669	−0.116	−0.088	−0.195
	(3.28)*	(3.29)*	(1.82)**	(3.02)*	(2.42)*	(2.00)*	(1.93)**	(2.21)*
2SLS								
1960–1983	0.781	−0.016	−0.082	−0.017	0.967	−0.022	−0.156	−0.107
	(2.56)*	(3.55)*	(2.29)*	(2.97)*	(2.86)*	(2.83)*	(2.83)*	(2.45)*
1984–2009	0.599	−0.020	−0.230	−0.104	0.760	−0.156	−0.354	−0.178
	(2.25)*	(1.79)**	(3.12)*	(3.28)*	(2.42)*	(4.39)*	(1.77)**	(2.04)*
1960–1997	0.874	−0.001	−0.014	−0.008	0.733	−0.002	−0.008	−0.026
	(2.00)*	(1.68)**	(2.21)*	(2.37)*	(1.77)**	(3.44)*	(1.70)**	(2.83)*
1998–2009	0.712	−0.088	−0.022	−0.036	0.623	−0.029	−0.120	−0.195
	(2.37)*	(2.34)*	(1.83)**	(1.80)**	(2.69)*	(1.67)**	(1.86)**	(2.55)*

Notes: Absolute t-ratios are in parentheses below the coefficients. Aus and NZ means Australia and New Zealand.
* and ** indicate significance at 5 and 10% levels, respectively.

Stability tests

Finally, we assessed the stability of *M1* demand functions using the Aus (2) and NZ (1) estimated equations for whole- and sub-sample periods through application of Cumulative Sum (CUSUM) and CUSUMSQ and Nyblom (1989) type tests, as suggested in Bruggeman *et al.* (2003); note that the results of the stability tests for equation Aus (1) gave qualitatively similar results. To conserve space, we report only the CUSUMSQ (as shown in Figs 1–4) and Nyblom (Table 7) tests results for 1984–1998 and 1998–2009 sub-periods. The Nyblom test proposed by Bruggeman *et al.* (2003) uses score functions directly rather than their first-order Taylor expansions. These scores are computed for maximum (supremum) and average (mean) values over the period of analysis and we denote these tests as $Sup_{t\in T}\, Q_T^{(t)}(i)$ and $Mean_{t\in T}\, Q_T^{(t)}(i)$ where $i = S$. In small samples asymptotic p-values can yield misleading results and therefore we also report the bootstrap p-values. The null hypothesis is that parameters are stable (constant) over the period of analysis.

The results of CUSUMSQ and Nyblom stability tests illustrate that *M1* demand functions were unstable in both countries over the period 1984–1998, which may imply that the 1980s reforms did have a significant impact on the demand for money in both countries. However, this impact on stability was temporary, as stability of *M1* demand is not rejected after 1998. Further, *M1* stability is not rejected in the whole-sample period.

Table 6. Short run estimates, 1960–2009

	Aus (1)	Aus (2)	NZ (1)
Intercept	1.236	12.627	4.013
	(6.05)*	(7.81)*	(7.95)*
ECM_{t-1}	−0.246	−0.211	−0.113
	(6.06)*	(7.39)*	(6.20)*
$\Delta \ln M_{t-2}$	–	−1.267	–
		(2.36)*	
$\Delta \ln Y_{t-1}$	0.726	–	0.026
	(3.45)*		(2.44)*
$\Delta \ln Y_{t-2}$	–	1.266	–
		(6.48)*	
$\Delta \ln E_{t-2}$	–	−1.006	−3.200
		(3.41)*	(2.35)*
ΔR_{t-2}	−0.253	−0.677	–
	(4.59)*	(2.26)*	
Adjusted R^2	0.803	0.816	0.763
SEE	0.065	0.063	0.077
χ^2_{sc}	0.324	0.676	0.893
	(0.57)	(0.41)	(0.35)
χ^2_{ff}	3.325	3.063	0.259
	(0.17)	(0.38)	(0.61)
χ^2_{n}	0.371	0.500	1.085
	(0.83)	(0.78)	(0.58)
χ^2_{hs}	0.020	0.025	0.006
	(0.89)	(0.90)	(0.94)

Notes: Dependent variable in each regression is $\Delta \ln M_t$. Absolute *t*-ratios for the variables and the *p*-values for the chi-square (χ^2) tests are in parentheses. Aus and NZ signifies Australia and New Zealand.
*Indicates significance at the 5% level.

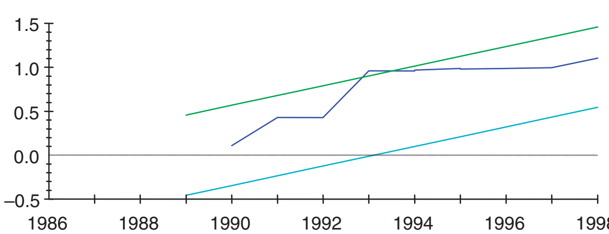

Fig. 1. Australian *M1* stability, 1984–1998 – plot of cumulative sum of squares of recursive residuals
Note: The straight lines represent critical bounds at 5% significance level.

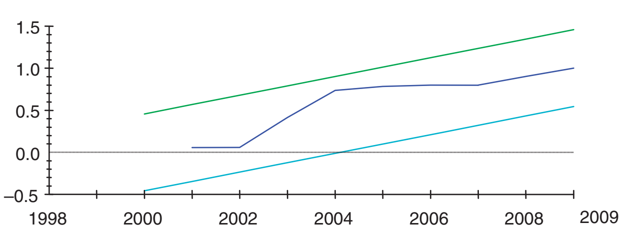

Fig. 2. Australian *M1* stability, 1998–2009 – plot of cumulative sum of squares of recursive residuals
Note: The straight lines represent critical bounds at 5% significance level.

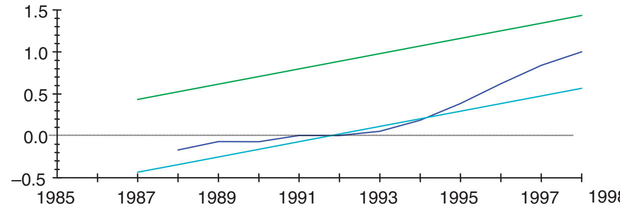

Fig. 3. New Zealand *M1* stability, 1984–1998 – plot of cumulative sum of squares of recursive residuals
Note: The straight lines represent critical bounds at 5% significance level.

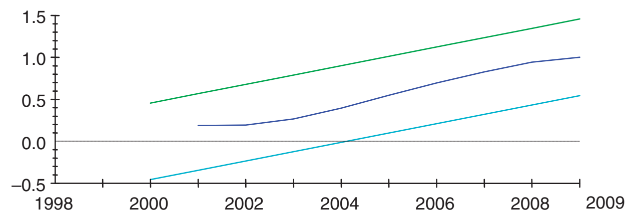

Fig. 4. New Zealand *M1* stability, 1998–2009 – plot of cumulative sum of squares of recursive residuals
Note: The straight lines represent critical bounds at 5% significance level.

The observed instability in money demand functions for both countries during the period 1984–1998 implies that it would have been appropriate monetary policy stance for their central banks to target the rate of interest. However, there is lack of evidence to support instability in the money demand functions after 1998, and therefore it would not be unreasonable if these central banks chose to switch policies and target the money supply as their instrument of monetary policy.

As emphasized by Poole (1970), the money supply (rate of interest) should be targeted if money demand is stable (unstable) and targeting the rate of interest when money demand is stable will accentuate instability in income. Under these circumstances, monetary targeting was the feasible policy stance for both countries.

V. Conclusion

This article has examined the demand for real narrow money (*M1*) for Australia and New Zealand over the period 1960–2009. Two specifications were considered: the canonical form and its extended form through augmentations of real effective exchange and inflation rates to capture the costs of

Table 7. Nyblom test for parameter stability

	$Sup_{t\in T}\,Q_T^{(t)}(S)$	p-value (Asym)	p-value (Boot)	$Mean_{t\in T}\,Q_T^{(t)}(S)$	p-value (Asym)	p-value (Boot)
Aus (2)						
1984–1998	3.260	0.000	0.000	8.041	0.000	0.030
1998–2009	0.587	0.946	0.548	0.322	0.985	0.870
NZ(1)						
1984–1998	1.152	0.032	0.058	2.542	0.000	0.014
1998–2009	0.199	0.801	0.569	0.456	0.357	0.625

Notes: p-values are asymptotic p-values. Comprehensive details on Nyblom stability tests are provided in Bruggeman *et al*. (2003).

holding money. Both specifications performed well for Australia but only the augmented version was plausible for New Zealand. The application of Lee and Strazicich's (2003) endogenous two break minimum LM unit root tests reveal that the variables (real *M1*, real income, nominal interest rate, real effective exchange rate and inflation rate) are $I(1)$ in levels.

Application of GH's method revealed that the cointegrating relationships of money demand underwent intercept and regime shifts in Australia and a regime shift in New Zealand. The results suggest a common break date of 1984; a break in the late 1990s was also present for both countries. Since the early 1980s both countries underwent continuing economic liberalization and the early break date may be capturing the circumstances of financial reforms. Estimates for the entire period reveal income elasticity estimates of around 0.64 and 0.68 for Australia and New Zealand, respectively, and the demand for money responds negatively to variations in the nominal rate of interest, and real effective exchange and inflation rates, albeit by small amounts.

Application of four time series methods viz., GETS, EG, FMOLS and 2SLS gave consistent results for two sets of sub-samples with 1984 and 1998 break dates. The income (interest rate) elasticities in both canonical and extended equations declined (increased) in the post-reform sub-samples. This illustrates improvements in the financial system around the break dates that are closely associated with the financial reforms.

Stability tests showed that money demand functions were unstable in the period 1984–1998 for both countries. The structural changes around 1984 did have a significant though temporary impact on the demand for money as the stability of *M1* demand is not rejected after 1998. These findings imply that it would not have been unreasonable for their central banks to use the rate of interest as an instrument of monetary policy during the period of instability and, following Poole (1970), monetary targeting when the money demand is stable.

Future research could examine the nature of financial reforms and their individual impacts on the demand for money. Given that a number of reforms have been implemented since the 1980s along with a number of other important events, it would be useful to analyse their impacts more specifically. Further research could use structural break tests to examine the stability of broad money for both countries.

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Appendix

Table A1. Johansen tests for cointegration, 1960–2009

	Canonical specification						Extended specification					
	Eigenvalue			Trace			Eigenvalue			Trace		
	Test statistic	95%	90%	Test statistic	95%	90%	Test statistic	95%	90%	Test statistic	95%	90%
Aus												
$r=0$	14.275	19.470	10.224	10.580	17.117	8.554	24.754	32.175	29.100	14.653	15.250	11.267
$r=1$	12.360	10.550	17.252	19.231	15.560	20.342	27.026	21.538	28.440	16.271	10.260	17.600
NZ												
$r=0$	12.034	11.930	10.672	23.260	18.250	14.360	23.717	20.480	19.270	45.236	36.248	35.140
$r=1$	8.270	12.425	10.280	27.872	31.280	29.050	11.230	14.980	13.645	14.260	15.248	19.230

Note: r is the number of cointegrating vectors.

Table A2. Johansen cointegration equations, 1960–2009

	Aus		NZ	
	Canonical specification	Extended specification	Canonical specification	Extended specification
$\ln y$	0.726	0.841	1.835	2.098
	(3.40)*	(2.75)*	(3.84)*	(1.87)**
R	−0.019	−0.279	−0.126	−0.105
	(1.76)**	(3.43)*	(2.32)*	(2.17)*
$\ln E$		−0.003		−0.028
		(1.47)		(2.01)*
π		−0.104		−0.073
		(2.08)*		(1.69)**

Notes: Absolute t-ratios are reported below the coefficients in parentheses.
* and ** indicate significance at the 5 and 10%, respectively.

Table A3. ADF unit root tests for residuals

	Canonical specification		Extended specification	
	Aus	NZ	Aus	NZ
1960–1983	−7.360 [1]	−4.548 [2]	−6.842 [0]	−8.300 [1]
	(−3.562)	(−3.567)	(−3.917)	(−3.439)
1984–2009	−4.552 [3]	−6.003 [0]	−3.970 [2]	−6.484 [0]
	(−3.567)	(−3.567)	(−2.879)	(−3.567)
1960–1997	−9.036 [2]	−5.265 [1]	−11.274 [1]	−4.404 [0]
	(−3.439)	(−3.917)	(−3.330)	(−3.917)
1998–2009	−10.271 [1]	−4.022 [1]	−5.720 [1]	−9.280 [2]
	(−2.879)	(−3.567)	(−3.562)	(−3.917)

Notes: The lag lengths are provided in square brackets. The ADF 5% critical values are given below the test statistics in parentheses. The lag lengths are chosen based on the following criteria:
(1) Set an upper bound $LAGmax$ for LAG.
(2) Estimate the ADF test regression with $LAG = LAGmax$.
(3) If the absolute value of the t-statistic for testing the significance of the last lagged difference is greater than 1.6 then set $LAG = LAGmax$ and perform the unit root test. Otherwise, reduce the lag length by one and repeat the process.