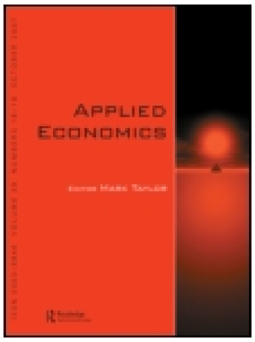

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Saten Kumar , Mamta B. Chowdhury & B. Bhaskara Rao

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Demand for money in the selected OECD countries: a time series panel data approach and structural breaks

Saten Kumar[a,*], Mamta B. Chowdhury[b] and B. Bhaskara Rao[b]

[a] *Department of Economics, Auckland University of Technology, Auckland, New Zealand*
[b] *School of Economics, University of Western Sydney, Sydney, NSW, Australia*

Time series panel data estimation methods are used to estimate the cointegrating equations for the demand for money (M1) for a panel of 11 Organization for Economic Cooperation and Development (OECD) countries for which consistent quarterly data are available. The effects of financial reforms are analysed with structural break tests and estimates for alternative sub-samples. Our results for the post-reform sub-samples show that the income elasticity of the demand for money has decreased and response to interest rate changes has increased.

Keywords: demand for money; Pedroni method; financial reforms; Westerlund method

JEL Classification: C50; E41

I. Introduction

Estimates of the demand for money and their stability have become controversial after 1970s due to the instabilities caused by financial reforms. Reforms have improved efficiency of the financial markets. Several money substitutes for transactions, e.g. credit and debit cards, electronic money transfers, etc., are created. Reforms have enhanced competition and also improved international capital mobility. It is now a stylized fact that the demand for various monetary aggregates has become unstable in the advanced countries following these reforms. Furthermore, developments in the estimation methods with the time series methods have raised doubts on the validity of the earlier estimates based on the classical methods and partial adjustment mechanisms (Taylor, 1994). Consequently, central banks in many advanced countries have switched from using money supply to the rate of interest as their instrument of monetary policy since it is not possible to forecast accurately the target with an unstable demand for money.[1] This is also consistent with Poole's (1970) analysis. According to him, money supply should be targeted when the demand for money is stable and the rate of interest when this relationship is unstable. Use of an incorrect instrument will only accentuate instability.[2]

Although stability of the demand for money in the advanced countries has been investigated by some previous studies, there do not seem to be many recent studies with the exception of Friedman and Kuttner (1992), Ball (2001), Nielsen (2004) and

*Corresponding author. E-mail: kumar_saten@yahoo.com

[1] Some studies support the Taylor rule-based interest rate targeting. See Orden and Fisher (1993) for Australia, McPhail (1991) and Haug (1999) for Canada, Nagayasu (2003) and Maki and Kitasaka (2006) for Japan, Papadopoulos and Zis (1997) for Greece, Vega (1995, 1998) for Spain, Oxley (1983) and Caporale and Gil-Alana (2005) for UK and Breuer and Lippert (1996) for USA.

[2] Following the advanced countries, central banks in many developing countries have also switched to targeting the rate of interest without any credible evidence that their money demand functions have become unstable. This shift has been questioned by Bahmani-Oskooee and Rehman (2005), Bahmani-Oskooee and Gelan (2009), Rao and Kumar (2009), Sumner (2009) and Yu and Gan (2009). These authors have found no instability in the demand for money functions using alternative estimation methods and data mainly from the Asian countries.

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Rao and Kumar (2011).[3] While Friedman and Kuttner and Ball found that the demand for money is unstable in the post-reform samples, Rao and Kumar (2011) found that the US demand for money has now become stable. Nielsen found similarly the demand for money in the UK is stable in the post-reform samples. As a plausible reason, it may be suggested that the improved stability may be due to wearing off the major effects of reforms by now. Given these findings, the contribution of this article is twofold. First, we shall use panel time series methods to estimate the demand for narrow money for 11 advanced Organization for Economic Cooperation and Development (OECD) countries, for which consistent quarterly data are available without gaps, and second, we test the stability of this function with the structural breaks test of Westerlund (2006). Since these OECD countries have implemented a number of financial reforms during much of the 1970s and 1980s, the effectiveness of such reforms may be observed as structural breaks in the cointegrating equations. If financial reforms were effective, there should be some increased economies of scale. Therefore it is expected that the income elasticity should show a decline and the response to changes in the rate of interest should increase. If the demand for money has not become stable by now, there should not be well-defined cointegrating equations for the whole sample period and also for the post-reform subsample periods. With these objectives in mind, the outline of this article can be stated as follows. Section II briefly reviews a few relevant empirical works. In Section III presents empirical results for panel unit root and cointegration tests and estimates of the cointegrating equations. Test results for structural breaks are presented in Section IV and Section V concludes this article.

II. Review of Panel Data Studies

Although many recent empirical studies on the demand for money have used country-specific data and time series methods, a number of studies have also used panel data and panel estimation methods (Mark and Sul, 2003; Valadkhani and Alauddin, 2003; Harb, 2004; Garcia-Hiernaux and Cerno, 2006; Lee and Chang, 2006; Dreger *et al.*, 2007; Elbadawi and Schmidt-Hebbel, 2007; Carrera, 2008; Hamori, 2008; Hamori and Hamori, 2008; Valadkhani, 2008; Fidrmuc, 2009; Rao and Kumar, 2009; Rao *et al.*, 2009; Setzer and Wolff, 2009).[4] Essentially, these studies, both on the developed and developing countries, have estimated one or another of the following specifications for the demand for money

$$\ln M_{it} = \alpha_1 + \alpha_2 \ln Y_{it} + \alpha_3 r_{it} + \varepsilon_{it} \tag{1}$$

$$\ln M_{it} = \alpha_1 + \alpha_2 \ln Y_{it} + \alpha_3 r_{it} + \alpha_4 \ln E_{it} + \varepsilon_{it} \tag{2}$$

$$\ln M_{it} = \alpha_1 + \alpha_2 \ln Y_{it} + \alpha_3 r_{it} + \alpha_4 \ln E_{it} + \alpha_5 \pi_{it} + \varepsilon_{it} \tag{3}$$

where M is real money stock, Y is real income, r is nominal rate of interest, E is real effective exchange rate, π is inflation rate, i and t are country and time subscripts, respectively, and $\varepsilon_{it} \sim N(0, \sigma)$ for all i and t.

Table 1 summarizes the estimated income elasticities and their main findings of some of these empirical works. While the majority of these panel data studies found that the income elasticity is near unity, estimates by Garcia-Hiernaux and Cerno (2006) are low at about 0.2 and that by Dreger *et al.* (2007) is high, exceeding 1.7. We shall treat these two studies as exceptional to the findings of the majority of the panel data studies. As stated before, it is to be expected that in the post-reform samples the income elasticity for the OECD countries is expected to be lower because of the improvements in the efficiency of the financial markets.

Before we present our estimates and examine the stability of the demand for money, it would be useful to briefly state a few details of the works in Table 1. The samples of Mark and Sul (2003), Dreger *et al.* (2007), Hamori and Hamori (2008), Fidrmuc (2009) and Setzer and Wolff (2009) are for varying number of the OECD countries. Mark and Sul (2003) have applied panel Dynamic Ordinary Least Squares (DOLS) method to estimate the demand for M1 using a panel of 19 OECD countries from 1957 to 1996. When a time trend was included, the estimated panel income elasticity was 1.08 with an interest rate semi-elasticity of –0.02. Fidrmuc (2009) used Pedroni's panel Fully Modified OLS (FMOLS) and Kao's (1999) panel DOLS techniques to examine M2 demand for six Central and Eastern European Countries (CEEC), namely, Czech Republic, Hungary, Poland, Romania, Slovakia and Slovenia over the period 1994 to 2003. They found that money demand depends significantly on the euro area interest rates and the exchange rate against the euro. Dreger *et al.* (2007) have used quarterly data for the period 1995(Q1) to 2004(Q2) and three panel cointegration techniques (Pedroni, 2000; Mark and Sul, 2003; Breitung, 2005) to estimate broad money (M3) demand for 10 EU countries.[5] The estimate of the panel income elasticity was between 1.73 and 1.94, while the interest rate elasticity is negative and significant. Dreger *et al.* (2007) argued that sudden introduction of the euro in all new EU member states may have introduced problems for the stability of the euro area money demand function. Hamori and Hamori (2008) have estimated the money demand (M1, M2 and M3) using the Pedroni FMOLS method and data from January 1999 to March 2006, covering the 11 EU countries.[6] They find that there is a long-run relationship between monetary aggregates and its determinants. Setzer and Wolff (2009) have estimated demand for M3 for euro area using the panel DOLS technique. They found that the income elasticity for M3 was around 1.67 for the sample 2003–2008, while for the sample starting from 2001, this was 1.20. Their findings imply that income elasticity has increased, which is contrary to

[3] Nielsen (2004) examined the UK money demand (M2, M3 and M4) for the period 1873 to 2001 and stability results suggest that the long-run demand for money is stable. According to Friedman and Kuttner (1992), the US demand for M1 is cointegrated for the period 1960 to 1979, but becomes unstable if the sample is extended from 1980. Ball's (2001) study for the US, notes that stability tests did not show breaks in the demand for M1 up to 1987, but becomes unstable when the samples are extended up to 1996.

[4] The cross-section and time series studies are well-surveyed earlier by Sriram (1999, 2001). Recently, Duca and VanHoose (2004) provided a useful survey of the theoretical and empirical literature on the demand for money.

[5] These EU countries are Cyprus, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Slovak Republic and Slovenia.

[6] These countries are Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, Netherlands, Portugal and Spain.

Table 1. Panel data studies on money demand and their findings

Author	Period; monetary aggregate	Country	Specification and methodology	Panel income elasticity	Panel interest rate elasticity	Main findings
Garcia-Hiernaux and Cerno (2006)	1988–1998; M0 (monetary base)	27 developed and developing countries	Equation 1; GMM	0.180 (6.80)* 0.200 (2.00)*	–0.004 (2.90)* –0.005 (1.70)**	The income elasticity is in contrast to Cagan's (1956) findings.
Dreger *et al.* (2007)	1995(Q1)–2004(Q2); M2	10 EU countries	Equation 2; Pedroni Mark and Sul Breitung	1.730 (21.63)* 1.940 (14.92)* 1.780 (17.80)*	–0.090 (4.50)* –0.070 (2.33)* –0.060 (3.00)*	Introduction of euro in new EU member states may have created instability in the euro area money demand.
Mark and Sul (2003)	1957–1996; M1	19 OECD countries	Equation 1; DOLS	1.079 (4.09)*	–0.022 (3.67)*	Without time trend, income elasticity is 0.860.
Harb (2004)	1979–2000; M1 and M2	6 Gulf Cooperation Council (GCC) countries	Equation 2; Pedroni	0.780 (11.48)* 0.420 (5.52)*	–0.050 (2.36)* 0.010 (0.33)	Real M1 yields more robust relation than real M2.
Carrera (2008)	1948–2003; M1	15 Latin-American countries	Equation 1; Pedroni	0.940 (50.20)*	–0.008 (11.43)*	The country-specific income elasticity is around unity in many Latin American countries.
Valadkhani and Alauddin (2003)	1979–1999; M2	8 Asian countries	Equation 3; SUR	n/a	n/a	The country-specific income elasticities range between 0.3 to 1.4.
Rao and Kumar (2009)	1970–2005; M1	14 Asian countries	Equation 1; Pedroni Mark and Sul Breitung	1.140 (20.84)* 0.990 (32.00)* 0.960 (60.19)*	–0.020 (5.60)* –0.010 (2.75)* –0.010 (5.24)*	Demand for M1 is stable in these countries.
Rao *et al.* (2009)	1970–2007; M1	11 Asian countries	Equation 3; SGMM	1.190 [0.00]* 1.160 [0.00]*	–0.543 [0.00]* –0.512 [0.00]*	There exists well-defined M1 demand with no structural breaks.
Setzer and Wolff (2009)	2003(Q1)–2008(Q3); M3	Euro area	Equation 1; DOLS	1.670 (5.90)*	–0.090 (3.12)*	Strong money growth has not altered demand for M3.
Hamori (2008)	1980–2005; M1 and M2	35 Sub-Saharan African countries	Equation 1; Pedroni	0.86 (32.79)* 1.00 (40.37)*	–0.020 (6.63)* –0.010 (2.76)*	Money supply is a reliable monetary policy instrument.

Notes: *t*-Statistics are given in the parentheses. In Rao *et al*. (2009) *p*-values are given in the square brackets. Valadkhani and Alauddin (2003) did not report income elasticity. DOLS, GMM, SGMM and SUR mean, respectively, Dynamic Ordinary Least Squares, Generalized Method of Moments, Systems Generalized Method of Moments and Seemingly Unrelated Regressions. In all cases, semi-interest rate elasticity is provided, except Rao *et al*. (2009).
* and ** denotes significance at the 5 and 10% levels, respectively.

Table 2. Panel unit root tests 1975(Q1)–2008(Q4)

Series	LLC	Breitung	IPS	ADF	PP	Hadri
ln M	0.375	0.488	1.672	9.247	8.253	8.781
	(0.646)	(0.687)	(0.953)	(0.980)	(0.990)	(0.000)*
ln Y	–2.379	3.320	–0.579	29.536	19.745	10.371
	(0.009)*	(0.999)	(0.281)	(0.078)**	(0.474)	(0.000)*
r	–2.556	–0.414	–0.187	24.500	10.552	10.303
	(0.005)*	(0.340)	(0.426)	(0.221)	(0.957)	(0.000)*
ln E	2.408	2.404	3.664	9.780	6.945	8.281
	(0.992)	(0.992)	(1.000)	(0.972)	(0.997)	(0.000)*
π	–4.843	–4.092	–5.055	62.395	50.775	6.584
	(0.000)*	(0.000)*	(0.000)*	(0.000)*	(0.000)*	(0.000)*
Δln M	–17.190	–8.061	–14.704	195.964	192.819	2.800
	(0.000)*	(0.000)*	(0.000)*	(0.000)*	(0.000)*	(0.003)*
Δln Y	–14.720	–7.439	–13.649	178.701	183.202	3.390
	(0.000)*	(0.000)*	(0.000)*	(0.000)*	(0.000)*	(0.000)*
Δr	–18.003	–12.419	–15.339	206.323	296.527	4.126
	(0.000)*	(0.000)*	(0.000)*	(0.000)*	(0.000)*	(0.000)*
Δln E	–9.037	–1.324	–8.133	104.134	98.690	5.041
	(0.000)*	(0.093)**	(0.000)*	(0.000)*	(0.000)*	(0.000)*
$\Delta\pi$	–15.164	–9.512	–15.132	218.088	761.325	8.069
	(0.000)*	(0.000)*	(0.000)*	(0.000)*	(0.000)*	(0.000)*

Notes: Probability values are reported in the parentheses. Baltagi (2005) and Pesaran and Breitung (2005) provide detailed discussion of these tests.
* and ** denote the rejection of the null at the 5 and 10% levels, respectively.

expectation because financial reforms will allow improved scale economies thereby reducing the income elasticity.

Utilizing the panel data techniques (Pedroni, 2000; Mark and Sul, 2003; Breitung, 2005), Rao and Kumar (2009) have estimated the demand for M1 for a panel of 14 Asian countries for the period 1970 to 2005. In all cases, the panel income elasticities were around unity. Their results showed that there is a well-defined demand for M1 in these countries. Similar conclusions are made by Rao *et al.* (2009) for 11 Asian countries using panel Systems Generalized Method of Moments (SGMM) technique and data from 1970 to 2007.

Combining observations across Latin-American countries for the period 1948 to 2003, Carrera (2008) has found, with the Pedroni panel FMOLS method, the income elasticity as 0.94 and the semi-interest elasticity as –0.008. Hamori (2008) has used data over 1980 to 2005 and the Pedroni panel FMOLS method to estimate the demand for M1 and M2 for 35 Sub-Saharan African countries. The income elasticities are found to be 0.9 and 1 for M1 and M2, respectively. The semi-interest rate elasticity is also significant with expected negative sign.

However, in all these studies there are no formal tests for structural breaks in the relationship, except in Rao *et al.* (2009), where they have used Mancini-Griffoli and Pauwels (2006) test for structural breaks for the estimates with *SGMM*. Given that a number of major financial reforms were implemented by many OECD countries to enhance the efficiency of the financial sector, it is likely that structural changes might have taken place in the demand for money. In this article we shall examine this aspect of the demand for money with a test procedure for structural breaks in the time series panel data developed by Westerlund (2006).

III. Unit Roots and Cointegration

We shall use quarterly data for 11 OECD countries for the period 1975(Q1) to 2008(Q4). These countries are Australia, Canada, Japan, Korea, Italy, Mexico, Norway, Spain, Sweden, Switzerland and the US. These countries and the data period are selected because there are no gaps in the data on the variables and therefore our data set is balanced.[7] Definitions of the variables and sources of data are given in the Appendix. The dependent variable is the real narrow money M1.

We first tested for the order of the variables, namely, ln M, ln Y, r, ln E and π using the panel unit root tests of Levin, Lin and Chu (LLC, 2002), Breitung (2000), Im, Pesaran and Shin (IPS, 2003), Augmented Dickey–Fuller (ADF) Fisher χ^2, Phillips–Perron (PP) Fisher χ^2 and Hadri (2000). The panel unit root test results are given in Table 2. These tests gave fairly unambiguous results for ln M and ln E. The LLC, IPS, ADF, PP and Breitung tests in which the null is that the

[7] The International Financial Statistics (IFS) and the World Bank database did not publish data on M1 after 1998 for some major OECD countries (e.g. France, Germany, Greece, Ireland, Netherlands and the UK) and this has constrained our selection of countries. Furthermore, for compatibility we have used the IFS and World Bank sources for data.

Table 3. Panel cointegration tests 1975(Q1)–2008(Q4)

Test statistic	Equation 1	Equation 2	Equation 3
Panel ν-statistic	−0.777	−1.528	−1.782**
Panel σ-statistic	−1.196	2.100*	0.234
Panel $\rho\rho$-statistic	−1.696**	0.211	−0.488
Panel ADF-statistic	2.865*	1.760**	3.337*
Group σ-statistic	−6.483*	−1.798**	−2.344*
Group $\rho\rho$-statistic	−2.876*	−1.165	−1.934**
Group ADF-statistic	3.265*	1.802**	3.726*

Notes: The test statistics are distributed as $N(0, 1)$.
* and ** denote significance at the 5 and 10% levels, respectively.

Table 4. Estimates of the cointegration coefficients 1975(Q1)–2008(Q4)

	Equation 1	Equation 2	Equation 3
ln Y	0.870	0.803	0.825
	(40.20)*	(26.11)*	(30.39)*
r	–0.054	–0.020	–0.012
	(13.18)*	(9.37)*	(11.42)*
ln E		–0.021	–0.030
		(2.24)*	(2.98)*
π			–0.808
			(0.43)

Notes: t-ratios are given in the parentheses. These are FMOLS estimates without common trends. Since these are fixed effects estimates, intercepts are not reported.
*Denotes significance at the 5% level.

variable is nonstationary is not rejected at the 5% level. In the Hadri test with the null that the variable is stationary is also rejected for these two variables at the 5% level. For ln Y and r, all the tests show that they are nonstationary variables at the 5% level, except in the LLC test at the 1% level. For π all the tests show that it is stationary, except the Hadri test. With the exception of the Hadri test, all other tests show that the first differences of all the variables are stationary. Therefore, it is reasonable to infer that these variables are $I(1)$ in levels, except π. Since the Hadri test provides support that π is nonstationary, we proceed to test and estimate the cointegrating vectors for the specifications in Equations 1–3.

Since inflation rate is computed as the change in the Gross Domestic Product (GDP) deflator and the fact that first-differenced variables in macroeconomics data series are usually $I(0)$, it is not unreasonable that majority of the unit root tests support the inflation rate as stationary. A number of studies, using time series methods, have used the inflation rate as a proxy for cost of holding money in addition to the rate of interest (e.g. see Choi and Oxley (2004) for New Zealand, Juselius (1998) for Denmark, Bahmani-Oskooee and Shabsigh (1996) for Japan, Drake and Chrystal (1994) for UK, Baba *et al.* (1992) for USA and Orden and Fisher (1993) for Australia and New Zealand).

The results for cointegration tests for Equations 1–3 are given in Table 3. The majority of the reported seven tests show that there is cointegration between the variables at the 5% level, except for Equation 2. Only the panel ν and panel σ test statistics in Equation 1 and panel σ and panel $\rho\rho$ test statistics in Equation 3 are insignificant at the 10% level. It is well known that the two ADF tests have more power against the null and they reject the null of no cointegration at the 5% level. Therefore, it can be concluded that the variables in Equations 1 and 3 are cointegrated and a long-run demand for money function exists for the group as a whole and the members of the panel. For Equation 2, weak cointegration exists because amongst the seven tests only panel σ is significant at the 5% level and three of them (panel ADF, group σ and group ADF statistics) are significant at only 10% level.

Table 4 provides the estimated panel group cointegrating parameters with the Pedroni FMOLS method.[8] Estimates of income elasticity and semi-interest rate elasticity in Equations 1–3 differ only marginally and both are significant at the 5% level. The income elasticities are 0.870 and 0.825, respectively, and the coefficient of the rate of interest has the expected negative sign but differ because additional variables are used, besides the rate of interest, in Equations 2 and 3 to proxy the cost of holding money. In Equation 2, the exchange rate elasticity is significant at the 5% level. However, in Equation 3 the coefficient of exchange rate is significant at the 5% level but the rate of inflation is insignificant.[9] On the basis of the above estimates we may conclude that for the whole sample period, estimates of income elasticity are slightly lower than unity and money demand is responsive to changes in the cost of holding money.

IV. Impact of Financial Reforms

Controversy remains in the literature with the stability of money demand functions in the advanced countries. Most advanced countries are operating under inflation targeting regimes through which the short-term rates of interest are adjusted to attain price and macroeconomic stability. Evidence suggests that the switch from a focus on the money supply to the interest rate as the primary monetary policy instrument has taken place in advanced countries because of an instability in money demand functions (McPhail, 1991; Haug, 1999; Caporale and Gil-Alana, 2005; Maki and Kitasaka, 2006). This instability was largely caused by financial reforms and liberalization policies that were implemented from the late 1960s to early 1970s. However, it is likely that the effects of reforms are worn out by now and consequently money demand stability may have been reverted. To this end, targeting of the rate of interest when the money demand function is stable is inappropriate (Poole, 1970). This aspect of

[8] Estimates of the individual country cointegrating parameters are not reported to conserve space but are available from authors upon request. FMOLS estimates are without common trends. DOLS estimates are similar and not reported to conserve space.
[9] Deleting the inflation rate did not make any difference to the coefficients of other variables. However, deleting this variable made the cointegration tests weaker. Therefore, we retained this variable and proceed to test and estimate the cointegrating equations for sub-samples for Equations 1 and 3 only.

Table 5. Westerlund tests for structural breaks 1975(Q1)–2008(Q4)

Country	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
Break date	1987 (Q3)	1994 (Q2)	1983 (Q2)	1994 (Q2)	1991 (Q4)	1983 (Q1)	1987 (Q1)	1984 (Q3)	1988 (Q1)	1996 (Q3)	1986 (Q1)

Notes: Single endogenous break dates are reported. Countries are numbered from (1) to (11), for example (1) Australia, (2) Canada, (3) Japan, (4) Korea, (5) Italy, (6) Mexico, (7) Norway, (8) Spain, (9) Sweden, (10) Switzerland and (11) US. Westerlund tests with DOLS option gave the same break dates.

money demand stability requires empirical investigation and our article attempts to fill this gap.

We shall examine the effects of the financial reforms on the demand for money in our sample of the 11 OECD countries. Financial reforms have been implemented in most of these OECD countries and it is difficult to argue that all these countries have undergone reform process at the same time. Indeed, it is difficult to select a common break date for all the countries in our panel. Therefore we utilize the recently developed Westerlund (2006) structural break tests to investigate the break dates in our sample. Table 5 reports the single endogenous break results in the intercepts and trends.[10]

It can be observed explicitly that these countries have undergone structural changes which we assume is due to financial reforms at different, but close time periods. US is the dominant OECD country with a break date at 1986(Q1). There are seven countries (Australia, Japan, Mexico, Norway, Spain, Sweden and US) with early break dates somewhat during 1980s and close to the break date for the US. However, four countries (Canada, Korea, Italy and Switzerland) have delayed break dates during 1990s, which may be due to a late start of financial reforms. Although it is difficult to develop sub-samples based on different break dates, we adopt the following pragmatic approach to analyse the impact of financial reforms. First, we use the US break date of 1986(Q1) because it is the dominant country in our sample. In this case, our sub-sample periods are 1975(Q1)–1985(Q4) and 1986(Q1)–2008(Q4). Second, we group the countries into two types, i.e. countries that had an early break date and those that had a late break date. It is worth noting that US is one of the seven countries that have an early break date, and therefore it would be reasonable to construct the sub-samples for these countries as 1975(Q1)–1985(Q4) and 1986(Q1)–2008(Q4). Conversely, Canada and Korea are amongst the four countries that had a common late break date, i.e. 1994(Q2). To this end, we select the sub-samples as 1975(Q1)–1994(Q1) and 1994(Q2)–2008(Q4) for these four countries.

Before further discussion, it would be useful to take an overview of what is expected from these sub-sample estimates. If financial reforms have been effective, it is to be expected that there will be evidence for some economies of scale in the use of M1 and also the response to the rate of interest will improve because of more market-based interest rate policies. Consequently, it is to be expected in the second set of sub-samples that income elasticity will show a decline and the absolute value of the interest rate coefficient will increase. The instability in the demand for money may also be observed if reforms have generated considerable quantity of near monies. This should be reflected in the second set of sub-samples as lack of a well-defined long-run relationship between money and its determinants, i.e. cointegration tests might show that there is no cointegration. The details for the cointegration tests for the sub-samples are reported in Table 6.

In the three sets of sub-samples, the null of no cointegration is rejected by the majority of the cointegration tests at the 5% level. The only exception is Equation 3 in the sub-sample 1994(Q2)–2008(Q4) where panel υ, panel σ, group σ and group $\rho\rho$-statistics are insignificant at the 5% level, however, group σ statistic is significant at the 10% level. The more powerful ADF test statistics are significant at the 5% level in all cases. Therefore this provides strong evidence that there is cointegration in the three sets of sub-sample periods.

The cointegrating coefficients for three sets of sub-samples, with alternative break dates, are given in Table 7. All the estimated coefficients have the expected signs and are significant at the 5% level, except for estimates of the coefficient of inflation rate in Equation 3. The estimates of Equation 3 do show that in the post-reform periods the income elasticity has declined and the rate of interest coefficient has increased in absolute value. Estimates of Equation 1 in first set of sub-sample (all countries) show that in post-reform period the income elasticity has decreased from 0.92 to 0.66 and the coefficient of rate of interest (absolute value) has increased from 0.007 to 0.075. Fairly similar change in elasticities is observed in the set of sub-samples for countries with the early and late breaks. For countries with early break income elasticity has declined from 0.81 to 0.70 and for the late break countries from 0.71 to 0.62. The absolute value of the coefficient of the rate of interest has increased from 0.004 to 0.098 and from 0.010 to 0.051 for the early and late break countries, respectively. These results imply that in the post-reform period scale economies have improved although this is more in the seven countries with an early break. Estimates of Equation 3 also support these conclusions although the magnitude of these effects is smaller. On this basis, we can conclude that reforms did improve the expected scale and rate of interest effects. There is also no evidence that a well-defined long-run demand for money function does not exist.

[10] Although it is also possible to test for multiple breaks, we decided to test for one dominant break because our data covers only a shorter period of about 30 years and after some immediate and initial effects of reforms might have taken place, i.e. from 1975 due to gaps in data. Multiple breaks may also give conflicting break dates and increase the number of sub-samples. Therefore, testing for a single dominant break is a pragmatic option.

Table 6. Panel cointegration tests for the sub-samples

	Pre-reform sub-samples			Post-reform sub-samples		
	All countries	Early break countries	Late break countries	All countries	Early break countries	Late break countries
Test statistic	1975(Q1)–1985(Q4)	1975(Q1)–1985(Q4)	1975(Q1)–1994(Q1)	1986(Q1)–2008(Q4)	1986(Q1)–2008(Q4)	1994(Q2)–2008(Q4)
Panel υ-statistic						
Equation 1	1.677**	2.079*	1.899**	–2.459*	2.313*	–5.138*
Equation 3	–1.767**	–3.012*	1.039	–1.229	2.060*	–0.622
Panel σ-statistic						
Equation 1	–2.916*	2.103*	–1.406	–1.685**	–4.712*	–3.924*
Equation 3	2.088*	3.416*	–3.033*	–2.394*	–0.263	–0.983
Panel $\rho\rho$-statistic						
Equation 1	–4.005*	–3.778*	–6.190*	–1.117	–1.034	–1.115
Equation 3	–0.552	2.076*	–3.694*	–1.702**	–3.699*	–4.117*
Panel ADF-statistic						
Equation 1	3.550*	4.081*	–4.427*	2.113*	3.817*	2.034*
Equation 3	2.189*	2.237*	–2.413*	2.675*	3.762*	2.006*
Group σ-statistic						
Equation 1	–2.423*	–2.059*	–0.531	–3.280*	–3.266*	–5.307*
Equation 3	2.382*	2.474*	–1.987*	–3.352*	2.748*	–1.856**
Group $\rho\rho$-statistic						
Equation 1	–3.046*	–1.305	–5.981*	–2.395*	–2.607*	–0.755
Equation 3	–1.313	–0.855	–2.854*	–2.221*	–0.286	–1.542
Group ADF-statistic						
Equation 1	2.117*	3.775*	–5.189*	2.283*	2.878*	2.086*
Equation 3	1.988*	3.009*	–4.180*	3.294*	–2.200*	3.861*

Notes: Australia, Japan, Mexico, Norway, Spain, Sweden and US are the countries that have an early break. Countries that have a late break are Canada, Korea, Italy and Switzerland. The test statistics are distributed as $N(0,1)$. The critical values at 5 and 10% levels are 1.96 and 1.64, respectively.
* and ** denote significance at the 5 and 10% levels, respectively.

Table 7. Estimates for sub-period cointegration coefficients

		$\ln Y$		r		$\ln E$	π
	Sub-sample periods	Equation 1	Equation 3	Equation 1	Equation 3	Equation 3	Equation 3
Dependent variable: log(M1)							
All countries	1975(Q1)–1985(Q4)	0.916	1.011	–0.007	–0.008	–0.166	–0.134
		(21.00)*	(14.98)*	(14.33)*	(13.33)*	(2.26)*	(1.27)
	1986(Q1)–2008(Q4)	0.656	0.785	–0.075	–0.052	–0.155	–0.490
		(23.20)*	(19.72)*	(11.75)*	(10.04)*	(1.48)	(0.47)
Early break countries	1975(Q1)–1985(Q4)	0.805	0.766	–0.004	–0.007	–0.304	–0.141
		(14.28)*	(12.44)*	(12.19)*	(10.96)*	(3.80)*	(0.24)
	1986(Q1)–2008(Q4)	0.700	0.611	–0.098	–0.059	–0.085	–0.965
		(13.81)*	(10.61)*	(9.59)*	(6.52)*	(1.68)**	(1.77)**
Late break countries	1975(Q1)–1994(Q1)	0.705	0.799	–0.010	–0.009	–0.062	–0.341
		(33.72)*	(24.73)*	(7.41)*	(6.85)*	(1.65)**	(0.82)
	1994(Q2)–2008(Q4)	0.616	0.715	–0.051	–0.041	–0.013	–0.276
		(15.81)*	(12.04)*	(6.79)*	(7.94)*	(3.10)*	(1.54)

Notes: Australia, Japan, Mexico, Norway, Spain, Sweden and USA are the countries that have an early break. Countries that have a late break are Canada, Korea, Italy and Switzerland. These are FMOLS estimates without common trends. DOLS estimates gave similar results and not reported to conserve space. The t-ratios are given in the parentheses.
* and ** denote significance at the 5 and 10% levels, respectively.

We attained income elasticities of M1 demand below unity and this is consistent with the Baumol (1952) and Tobin (1956) model. Baumol–Tobin model analyse the costs and benefits of holding money from transactions perspective and explains how the use of money in absolutely foreseen transactions implies economies of scale and offers an interest rate elasticity that is significantly different from zero (Serletis, 2001, p. 78). Westerlund's structural break results imply that there has been an intercept and trend shift in all the selected countries and it is likely that holding of M1 balances may have reduced. This inference is valid because we observed scale economies in M1 demand, which implies that holdings of M1 balances may have declined. To this end, it is not necessary to estimate an inverse of velocity function.[11]

We assumed that the structural breaks provided by the Westerlund (2006) tests signify the phases of financial reforms and if these reforms have been effective and have led to any instability in the demand for money, then this would be observed in the post-reform sub-samples as a lack of a well-defined long-run relationship between money and its determinants. It is also possible that the variables in the post-reform samples change from being nonstationary to stationary, implying that there exists no cointegration. However, we found that all variables are $I(1)$ in levels in both sub-samples (except the inflation rate)[12] and there exists cointegration among the variables. In most cases, the income (rate of interest) elasticity has declined (increased) in the post-reform sub-samples. Our results imply that financial reforms have increased the scale economies of M1 demand but were not effective to cause any instability in the M1 demand. There seems to be no significant evidence that the M1 demand in the developed countries has become unstable. Nevertheless, central banks in many developed countries have switched to the bank rate as their instrument of monetary policy. Such an inconsiderate choice of monetary policy instrument could actually lead to increased instability in the output. Hence, following Poole (1970), we infer that the supply of money appears to be the feasible monetary policy instrument to be utilized by the central banks of advanced countries.

V. Conclusion

This article has used time series panel data technique of Pedroni (panel FMOLS) to estimate the long-run demand for money (M1) for a panel of 11 OECD countries. Estimates for the entire sample period of 1975(Q1) to 2008(Q4) showed that income elasticity of demand for money is lower than unity and demand for money responds negatively to variations in the rate of interest. We tested if the financial reforms in these countries had any significant effects. Our sub-sample estimates show that reforms have reduced the income elasticities and the rate of interest of semi-elasticity has increased. In the context of money demand, this highlights improved economies of scale, payment technology and the use of money substitutes.

An implication of our results is that, financial reforms may have contributed to some instability in the demand for money. But when structural changes are allowed, the pre- and post-reform sub-sample estimates imply that there is a stable and well-defined demand function for money in both sub-samples. The changes in the estimated effects on the parameters seem to be marginally higher in the countries that have implemented the reforms early. Another implication of our findings is that the central banks in these countries should reconsider their choice of using the interest rate as their monetary policy instrument because, according to Poole, money supply should be used as the monetary policy instrument when the demand for money is stable. Recent shift towards quantitative targets by the central banks in Europe may be based on the realization that the demand for money may have been stable now because many major effects of reforms may have already taken place.

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[11] We thank a referee of this journal for asking us to make this point clear.

[12] Hadri tests provide support that inflation rate is $I(1)$; the sub-sample unit root tests are not reported to conserve space.

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Appendix: Data

- Y Real GDP at factor cost. Data are from IFS-2009.
- R The average of 1–3 years savings deposit rate. Data are from IFS-2009.
- M Real narrow money supply. Data are from IFS-2009.
- E Real effective exchange rate. Data are from IFS-2009.
- π Rate of inflation calculated with GDP deflator. Data are from IFS-2009.

Note: All real variables are their nominal values deflated with the GDP deflator.