Long-run Money Demand in OECD Countries: Cross-Member Cointegration

Abstract. This paper examines the long-run money demand function for 11 OECD countries from 1983 to 2006 using panel data and including wealth. The distinction between common factors and idiosyncratic components using principal component analysis allows to detect cross-member cointegration and to distinguish between international and national developments as drivers of the long-run relation between money and its determinants. Indeed, cointegration between the common factors of the underlying variables, i.e. cross-member cointegration, indicates that the long-run relationship is mainly driven by international stochastic trends. Furthermore, it is found that the impact of income and the exchange rate on money demand is positive, while it is negative for the interest rate and stock prices. The estimated (semi-)elasticities of money are larger for the common components than for the original variables, except the income elasticity. Finally, the results of a panel-based error-correction model suggest that money demand converges to an international cross-member equilibrium relation of the common components.

JEL-Classification: E41, C22, C33

Keywords: Money demand, wealth effects, panel unit roots, panel cointegration, vector error-correction models

1 Introduction

The stability of the long-run money demand function is a widely studied research topic. On the one hand, stable money demand is relevant for policy makers to choose a sensible monetary policy instrument. For instance, unstable money demand caused by the financial reforms of the late 1970s induced many central banks in developed countries to switch from money targeting to the interest rate as monetary policy instrument. The very same was proposed by Poole (1970) who showed that the interest rate should be targeted if the money demand function is unstable. The identification of the optimal monetary policy strategy has also been studied for the upcoming European System of Central Banks (ESCB) in 1999. In this regard, many time series studies discussed the question whether monetary targeting or inflation targeting would be better to achieve price stability in the European Economic and Monetary Union (EMU). On the other hand, money can play an important role in the formulation of an efficient monetary policy strategy, even though monetary policy of developed countries typically uses an interest rate as policy instrument. Monetary aggregates can be appropriate indicators for future inflation in the medium term and long-run as long as there is a stable money demand function as it is also required for monetary targeting. As mentioned by Valadkhani (2006) an emerging consensus among economists is that it is not advisable to concentrate exclusively on a single policy instrument while neglecting another important information variable. Both the interest rate and monetary aggregates are important in selecting appropriate monetary policy actions. Monetary aggregates, however, will only be related to the real economy if the money demand function is stable. The stability of money demand entails whether money is an appropriate guide to policy.

Referring to that, the main purpose of this panel data study is to determine the existence of a stable long-run money demand function of 11 OECD countries from 1983 to 2006, taking into consideration possible cross-sectional dependencies resulting from common factors. If cross-member cointegration is existent the non-stationarity in the variables will stem from cross-sectional common stochastic trends only and the variables will be cointegrated across the panel members. To detect cross-member cointegration the common factors of the underlying variables are tested for unit roots and cointegration relations. The distinction between common factors and idiosyncratic components using principal component analysis also allows to distinguish between international and national developments as drivers of the long-run relation between money and its determinants (see Belke, Dreger, and Dobnik (2010)). Given that the idiosyncratic component is a residual, which captures the impact of shocks affecting the respective variable of one specific country, it can be interpreted as the part of the variable that is driven by national trends. In contrast, the common component represents international trends in the evolution of the variable, because it depends on a small number of common shocks, which affect the respective variable of all the countries. Depending on the results of the cointegration tests, this distinction has important implications for policy makers. If the common factors cointegrate, i.e. in the case of cross-member cointegration, the national monetary policy should take into account international developments, for instance, to precisely predict national future inflation. Indeed, this paper delivers empirical evidence that money and its determinants are cointegrated in their common factors, but not in their idiosyncratic components.

Moreover, this panel data study incorporates wealth as additional determinant of money demand as has become popular in country-specific time series studies. Friedman (1970) already suggested in 1970 that wealth may play an important role for money demand if it was viewed in a portfolio framework. Following Friedman (1988) and Choudhry (1996), among others, this paper introduces stock prices as wealth variable assuming that equities have a strong relation with money. In addition, the standard money demand function with income as scale variable and an interest rate as measure of opportunity cost is further extended by the exchange rate capturing possible currency substitution

effects. Furthermore, this paper contributes in applying an error-correction specification using panel data to determine long-run as well as short-run coefficients of the money demand of OECD countries. Panel data studies usually estimate only the long-run relation ignoring the short-run dynamics except the studies by Valadkhani (2008) and Nautz and Rondorf (2010).
The remainder of this paper is organised as follows. Section 2 briefly reviews the literature on money demand using panel data. Section 3 introduces the money demand function. Section 4 presents the data, discusses the econometric methods and presents the empirical results. Section 5 provides conclusions.

2 Review of panel data studies

While there is a wide range of country-specific time series studies on money demand available, only a few studies apply panel-econometric methods so far (see Table 1). The use of panel datasets provides more powerful unit root and cointegration tests compared to standard time series tests. It is widely known that standard unit root and cointegration tests based on individual time series have low statistical power, especially when the time series is short (Campbell and Perron, 1991). Panel-based tests rely on a broader information set by extending the time series dimension by the cross-sectional dimension, allowing for higher degrees of freedom. Therefore, the statistical power can substantially be increased and tests are more accurate and reliable. Table 1 summarises the studies using panel datasets to analyse the long-run relationship between money and its main determinants. The estimated income elasticities vary between 0.18 (Garcia-Hiernaux and Cerno, 2006) and 2.66 (Hamori and Hamori, 2008), but are usually slightly greater than one. The estimated interest rate (semi-)elasticities are in the range of -0.71 (Nautz and Rondorf, 2010) and 0.008 (Arnold and Roelands, 2010), where the latter value should be treated as an exception with a sign contrary to theory. For the exchange rate there exists no clear sign which can be imposed from theory. The few panel data studies including exchange rates come up with ambiguous results. The coefficients take values between -1.73 (Rao et al., 2009) and +0.31 (Narayan et al., 2009). Some of these studies use further additional explanatory variables for money beyond income, the interest rate and the exchange rate, among them inflation or a foreign interest rate. Time series studies, however, meanwhile often employ wealth as additional explanatory variable (see, e.g. Setzer and Greiber (2007); Boone and van den Noord (2008); de Santis et al. (2008); de Bondt (2009); Dreger and Wolters (2010))[1], but panel data studies usually do not. The studies by Arnold and Roelands (2010) and Nautz and Rondorf (2010)) are the only exceptions. According to Friedman (1988), stock prices as financial wealth variable may have two kinds of impacts on money demand, a positive wealth effect and a negative substitution effect. A wealth effect occurs in three different scenarios. First, a rise in stock prices leads to additional wealth which may be stored in money. Second, an increase in stock prices reflects an increase in the expected return from risky assets relative to safe assets. The resulting increase in relative risk may induce economic agents with given risk aversion/preference to hold larger amounts of safer assets such as money in their portfolio. Third, a higher level of stock prices may imply a rise in the volume of financial transactions, resulting in an increase in money demand to facilitate these transactions. In contrast, the negative substitution effect suggests that a rise in asset prices reduces the attractiveness

[1]Many time series studies on euro area money demand (e.g. Boone and van den Noord (2008)) argue in favour of the inclusion of wealth to explain the overshooting of the ECB's M3 target and to reestablish a stable money demand function. What is more, some studies use a (financial) wealth variable in addition to an income variable to take account of an income elasticity greater than one due to an omitted variables bias (see recent surveys by Knell and Stix (2005, 2006).

of holding money as a component of the portfolio. Consequently, the net impact of wealth on money demand has to be determined empirically. The study by Arnold and Roelands (2010) found a positive impact of house prices on money demand for the whole panel of ten euro area countries, but no significant impact of stock prices. Nautz and Rondorf (2010) cannot reject empirically that neither house prices nor stock prices significantly affect euro area long-run money demand. Further, most panel data studies estimated only the long-run coefficients of money demand. Valadkhani (2008) and Nautz and Rondorf (2010) also accounted for the short-run dynamics of money demand by estimating a panel-based error-correction model. The omission of a dynamic error-correction model to capture short-run dynamics of money demand might be due to the problem that there exist no tests to detect instability of panel estimated regression equations which correspond to the popular CUSUM and CUSUMSQ tests in country-specific time series models (see Rao and Kumar (2009)).[2] For a more detailed description of panel data studies on money demand including their methods and main findings see Kumar, Chowdhury, and Rao (2010).

Table 1: Overview of panel data studies on money demand

Study	Countries	M	Income Elasticity	Interest Rate (Semi-)Elasticity	Exchange Rate Elasticity
Mark and Sul (2003)	19 OECD count.	M1	1.08	-0.02	
Valadkhani and Alauddin (2003)	8 developing c.	M2	n/a	n/a	n/a
Harb (2004)	6 GCC count.	M1	0.78	-0.05	0.04
Garcia-Hiernaux and Cerno (2006)	27 countries	M0	0.18 to 0.20	-0.005 to -0.004	
Dreger et al. (2007)	10 EU count.	M2	1.73 to 1.94	-0.09 to -0.06	-0.28 to -0.16
Elbadawi and Schmidt-Hebbel (2007)	99 countries	M1	0.61 to 0.86	-1.13	
Hamori (2008)	35 Sub-Saharan	M1	0.86 to 0.89	-0.38 to -0.02	
	African count.	M2	1.00 to 1.02	-0.28 to -0.01	
Hamori and Hamori (2008)	11 EU count.	M1	2.52 to 2.66	-0.25 to -0.08	
		M2	1.50 to 1.59	-0.16 to -0.05	
		M3	1.73 to 1.82	-0.17 to -0.05	
Valadkhani (2008)	6 Asia-Pacific c.	M2	1.48	-0.03 to -0.02	-0.26 to -0.12
Fidrmuc (2009)	6 CEECs	M2	0.23 to 1.06	-0.009 to -0.002	-0.07 to -0.03
Narayan et al. (2009)	5 South Asian c.	M2	1.23 to 1.31	-0.23 to -0.20	0.26 to 0.31
Rao and Kumar (2009)	14 Asian count.	M1	0.94 to 1.14	-0.02 to -0.01	
Rao et al. (2009)	11 Asian count.	M1	0.94 to 1.98	-0.54 to -0.51	-1.73 to -0.87
Setzer and Wolff (2009)	Euro Area	M3	1.67	-0.09	
Arnold and Roelands (2010)	Euro Area	M3	1.55 to 2.60	-0.011 to 0.008	
Kumar et al. (2010)	11 OECD count.	M1	0.83 to 0.87	-0.05 to -0.01	-0.03
Kumar (2010)	5 Pacific Island c.	M1	0.98 to 1.06	-0.02 to -0.03	
Nautz and Rondorf (2010)	Euro Area	M3	1.41 to 1.55	-0.71 to -0.40	

[2]Stability tests like the CUSUM and CUSUMSQ are applied in this context to test the stability of the short-run coefficients and the adjustment coefficient of the lagged error-correction term. To test for the stability of the long-run money demand first the cointegration relation has to be estimated. In a second step, CUSUM and CUSUMSQ tests may be applied to test its stability.

3 Money demand function

In this paper a widely used specification of the money demand function is chosen as starting point. According to Ericsson (1998), the main body of theories of money demand assumes a long-run money demand function

$$\frac{M}{P} = f(Y, OC) \tag{1}$$

that relates real money balances M/P to a scale variable (Y), which represents real economic transactions, and the opportunity cost of holding money (OC), reflecting the forgone earnings due to holding alternative assets. M refers to a monetary aggregate in nominal terms and P denotes the price level. Presenting the money demand function in real terms of money implies that the demand for nominal money fully adjusts to price movements in the long run, so that the desired level of real balances remains unchanged. Hence, the use of real money balances as the dependent variable incorporates the assumption of long-run price homogeneity as predicted by most theories. Moreover, imposing a unitary price elasticity makes the identification problem between money demand and money supply less serious.[3] As in theoretical models, the empirical models also generally specify the money demand as a function of real money balances. In the empirical analysis a semi-logarithmic linear specification of long-run money demand is preferred. Panel data studies usually estimate one of the following specifications for money demand.

$$\ln M_{i,t} = \alpha_i + \beta_{1i} \ln Y_{i,t} + \beta_{2i} R_{i,t} + \varepsilon_{i,t}, \tag{2}$$
$$\ln M_{i,t} = \alpha_i + \beta_{1i} \ln Y_{i,t} + \beta_{2i} R_{i,t} + \beta_{3i} \ln EX_{i,t} + \varepsilon_{i,t}, \tag{3}$$
$$\ln M_{i,t} = \alpha_i + \beta_{1i} \ln Y_{i,t} + \beta_{2i} R_{i,t} + \beta_{3i} \ln EX_{i,t} + \beta_{4i} \pi_{i,t} + \varepsilon_{i,t}, \tag{4}$$

where $i = 1, ..., N$ represents panel members and $t = 1, ..., T$ denotes the time period. $M_{i,t}$ is the real money stock, $Y_{i,t}$ represents a measure of real income as a scale variable, $R_{i,t}$ is the nominal interest rate, $EX_{i,t}$ is the real exchange rate and $\pi_{i,t}$ stands for the inflation rate. The disturbance term $\varepsilon_{i,t}$ is assumed to be a white noise error process. Usually, real GDP represents the real income and, therefore, the transactions volume in the economy. The opportunity cost of holding money is proxied with the nominal interest rate and the inflation rate. The parameters of the models measure the (semi-)elasticities of money demand *vis-à-vis* the respective variables. The theoretically expected sign of the income elasticity of money demand, β_{1i}, is positive. More precisely, the quantity theory of money proposes a value of 1 for β_{1i} whereas the Baumol-Tobin model predicts a magnitude of 0.5 for β_{1i}. The interest rate and the inflation rate can be interpreted as rates of return that economic agents abandon by holding money instead of some alternative (financial or physical) assets. Consequently, the anticipated signs for the semi-elasticities for the interest rate and for the inflation rate are $\beta_{2i} < 0$ and $\beta_{4i} < 0$. The exchange rate is included with an eye on the literature on currency substitution which suggests that portfolio shifts between domestic and foreign money can be captured by the exchange rate. The expected sign of the elasticity of the exchange rate is less obvious. Any variation in the exchange rate can be argued to have both a positive and a negative impact on the demand for domestic currency. On the one hand, there is a positive currency substitution effect. A stronger

[3] The question may be raised, whether it is possible to estimate a money demand function without specifying simultaneously money supply. The problem can be avoided by assuming that money demand is independent of the price level. Since money supply is invariably specified in nominal terms across all competing theories, there exists no supply function for real balances and therefore no identification problem, see Laidler (1993).

domestic currency, i.e. an exchange rate appreciation, increases domestic money demand. On the other hand, a real exchange rate appreciation is also associated with a negative shock to economic activity and hence potentially also lowers domestic money demand.
In contrast to most other panel data studies, except Arnold and Roelands (2010) and Nautz and Rondorf (2010), this study additionally includes a wealth variable as it has become common in time series studies on money demand:

$$\ln M_{i,t} = \alpha_i + \beta_{1i} \ln Y_{i,t} + \beta_{2i} R_{i,t} + \beta_{3i} \ln EX_{i,t} + \beta_{4i} \ln W_{i,t} + \varepsilon_{i,t}, \tag{5}$$

where $W_{i,t}$ denotes wealth as a further determinant of money demand. According to Friedman (1988), stock prices as wealth variable may have both a negative substitution effect and a positive wealth effect as mentioned in section 2. Hence, the net impact of wealth on money demand has to be determined empirically.
This panel data study analyses the cointegration relation between money, income, the interest rate, the exchange rate and the stock prices in more precise terms. First, in order to detect cross-sectional dependencies in terms of cross-member cointegration and to distinguish between national and international trends as potential drivers of long-run money demand, each variable is separated into common and idiosyncratic components by a principal component analysis. Second, this study tests common factors and idiosyncratic components separately for unit roots and their cointegration properties. Third, the long-run (semi-)elasticities of money demand are estimated. As a final step, the short-run coefficients and the adjustment coefficient are determined using a panel error-correction model.

4 Data, methodology and empirical results

This study is based upon seasonally adjusted quarterly data from 1983 to 2006 for 11 OECD countries: Australia, Canada, Denmark, France, Germany, Italy, Japan, the Netherlands, Sweden, Switzerland, and the United States. The scale variable and the opportunity cost variable are represented by real GDP (Y) and the nominal three-month interbank rate (R), respectively. In line with the literature on currency substitution the benchmark money demand function including these variables is extended by the real effective exchange rate (EX). Further, following Friedman (1988) and Choudhry (1996), among others, real stock prices (W) are introduced as additional wealth variable. The dependent variable, real money (M), is the log-difference between the monetary aggregate M1 and the consumer price index (CPI). The use of a narrow monetary aggregate has several advantages. First, M1 is a good measure of liquidity in the economy since it consists mainly of financial assets held for transaction purposes. Second, the central bank is able to control this aggregate more accurately than broader aggregates such as M2 and M3. Third, M1 definitions tend to be relatively consistent across countries and, therefore, allow for comparison (Bruggeman, 2000). All variables are deflated with the CPI and expressed in natural logarithms, except the interest rate which is nominal and expressed in terms of levels. The CPI and the exchange rate have been obtained from the International Financial Statistics of the IMF. Data for monetary aggregates and interest rates have been taken from the Financial Indicators dataset of the OECD[4] and the GDP stems from the quarterly national accounts database of the OECD.
It is widely known that standard unit root and cointegration tests based on individual time series have

[4]The Financial Indicators dataset is a subset of the Main Economic Indicators (MEI) database of the OECD.

low statistical power, especially when the time series is short (Campbell and Perron, 1991). Panel-based tests represent an improvement in this respect by exploiting additional information that results from the inclusion of the cross-sectional dimension. However, most first generation panel unit root and cointegration tests assume that the cross-section members are independent. This condition is often likely to be violated, for example, because of common shocks to inflation such as those stemming from oil and food price increases. But most existing residual based tests use the assumption of cross-sectional independence to be able to get a nice asymptotic distribution for the test statistic. The independence of the cross-section members allows for the use of standard asymptotic tools, such as the Central Limit Theorem. Inappropriately assuming cross-sectional independence in presence of cross-member cointegration, however, can distort the panel results (see Banerjee et al. (2004), Urbain and Westerlund (2006)). Therefore, this study controls for cross-section dependencies by taking into account the common factor structure

$$Y_{i,t} = \xi_{1i} F_{1t} + E_{1i,t}, \quad \text{and} \tag{6}$$
$$X_{i,t} = \xi_{2i} F_{2t} + E_{2i,t}, \tag{7}$$

where F denotes the common factors and E stands for the idiosyncratic components of the respective variables. Gengenbach, Palm, and Urbain (2006) have proposed a sequential testing strategy based on the factor structure under equations (6) and (7) that does not restrict the heterogeneity and determines whether dependencies between the cross-sections are persistent. They consider two important cases. First, the common factors are I(1), while the idiosyncratic components are I(0). In this case non-stationarity in the panel is solely driven by a reduced number of common stochastic trends and cross-member cointegration may exist. A cointegration relationship between $Y_{i,t}$ and $X_{i,t}$

$$Y_{i,t} - \beta_i X_{i,t} = \xi_{1i}\left(F_{1t} - \beta_i \frac{\xi_{2i}}{\xi_{1i}} F_{2t}\right) + E_{1i,t} - \beta_i E_{2i,t} \tag{8}$$

can then occur only if the common factors of $Y_{i,t}$ cointegrate with those of $X_{i,t}$. The null hypothesis of no cointegration between these estimated factors can be investigated using standard time series tests such as the Johansen reduced rank approach (Johansen, 1995). The second case proposed by Gengenbach et al. (2006) denotes that both common and idiosyncratic stochastic trends are present in the data. Both the common factors and the idiosyncratic components are I(1) and have to be tested separately for cointegration. Cointegration between $Y_{i,t}$ and $X_{i,t}$ implies that both the common and idiosyncratic parts of the error term are stationary, see equation (8). Since the defactored series are independent by construction, cointegration between the idiosyncratic components can be explored by first generation panel cointegration tests such as those of Pedroni (1999, 2004). It should be noted, however, that the existence of cointegration relationships that annihilate both the common and idiosyncratic stochastic trends is very unlikely, see Gengenbach et al. (2006).

4.1 Variable decomposition

The first and innovative step of this paper regarding the long-run money demand function is to decompose each variable into the two uncorrelated components, i.e. a common and an idiosyncratic component, as suggested by Bai and Ng (2004). The idiosyncratic component is a residual, which captures the impact of shocks affecting the respective variable of one specific country. These country-specific shocks, such as domestic money demand shocks, may have large but geographically concentrated effects. The common component of a variable is 'common' in the sense that it depends on a small number of common shocks, which affect the respective variable of all the countries. The decomposition by principal component analysis is based on differenced data because of potential non-stationarity of the levels of the variables, as proposed by Bai and Ng (2004). After estimating the common factors they are re-cumulated to match the integration properties of the original variables. The idiosyncratic components are obtained from a regression of the original series on their common factors. For all variables two common components are enough to capture 50 to 70 percent of the overall variance. Any further component would add only a small proportion and the evidence shows that results do not qualitatively change.

As a second step, the common factors and idiosyncratic components are tested separately for unit roots and cointegration relationships. A cointegration relationship between the variables requires that the null hypothesis of no cointegration can be rejected for both the common and the idiosyncratic components, see equation (8). If the common factors are I(1), while the idiosyncratic components are I(0), the non-stationarity in the panel will be driven entirely by a reduced number of international stochastic trends and cross-member cointegration may exist.

4.2 Unit root tests

In the analysis of the common factors of real money, real GDP, the interest rate, the real effective exchange rate and real stock prices standard time series unit root tests can be applied. To test the null hypothesis of a unit root the augmented Dickey and Fuller (1979) (ADF) test and the Phillips and Perron (1988) (PP) test were used. According to the results displayed in Table 2 the common factors of money demand and its determinants all turn out to be non-stationary (with and without trend) and to become stationary by taking first differences. Hence, the results suggest evidence in favour of common factors that are integrated of order one, I(1).

Table 2: Time series unit root tests for common components

Variable	Levels				Differences	
	ADF	PP	ADF	PP	ADF	PP
	without trend		*with trend*			
M^c	-0.99(1)	-1.03[5]	-0.54(1)	-0.46[5]	-5.75(0)***	-5.78[3]***
Y^c	-1.97(9)	-1.10[5]	-2.94(9)	-2.12[5]	-5.12(0)***	-5.12[3]***
R^c	-2.34(9)	-1.67[5]	-3.06(9)	-2.01[5]	-4.32(4)***	-6.66[4]***
EX^c	-2.39(2)	-2.51[5]	-2.36(2)	-2.44[5]	-5.89(1)***	-8.34[4]***
W^c	-1.89(4)	-1.96[4]	-2.37(4)	-2.29[4]	-4.79(3)***	-8.35[4]***

Notes: The superscript *c* denotes the common factor of the respective variable. Numbers in parentheses are lag levels based on the Akaike Information Criterion. Numbers in brackets represents the automatic Newey-West bandwidth selection using the Bartlett kernel. ***, ** and * indicate significance at the 1%, 5% and 10% levels.

Stochastic trends in the idiosyncratic components can be efficiently explored by first generation panel unit root tests, since the defactored series are independent by construction and, thus, fulfil the assumption of cross-sectional independence. This study applies the Levin, Lin, and Chu (2002) (LLC) test, the Fisher-type ADF test and the Fisher-type PP test (see Maddala and Wu (1999) and Choi (2001)).

The LLC test restrictively assumes that all cross-sections have the same first order autoregressive parameter. By contrast, the non-parametric Fisher-type tests relax this assumption by allowing heterogeneity in this coefficient for all cross-section units. Although the Fisher-type tests are preferable according to that, this study also reports the results of the LLC test to provide an additional check for robustness.[5] In contrast to the time series unit root evidence for the common components, the panel unit root tests propose that the idiosyncratic components of the variables under investigation are widely stationary (see Table 3).

Table 3: Panel unit root tests for the idiosyncratic components

Variable	LLC	ADF-Fisher	PP-Fisher
M^i	-3.07***	36.46**	38.71**
Y^i	-3.10***	36.89**	32.40*
R^i	-5.65***	62.08***	62.80***
EX^i	-3.15***	34.36**	25.93
W^i	-4.45***	54.71***	50.30***

Notes: The superscript i denotes the idiosyncratic component of the respective variable. Probabilities for the Fisher tests are computed using an asymptotic Chi-square distribution. The LLC test assumes asymptotic normality. The choice of lag levels for the Fisher-ADF test is based on the Akaike Information Criterion. The LLC and Fisher-PP tests were computed using the Bartlett kernel with automatic bandwidth selection. ** and * indicate significance at the 1% and 5% levels.

Hence, the results indicate that random walks in the data are driven mainly by common international developments. In other words, the non-stationarity in real money and its determinants of the 11 OECD countries stems from common rather than country-specific shocks. As a consequence, a long-run equilibrium relationship for money demand may exist between the common rather than the idiosyncratic components, which would be equivalent with cross-member cointegration.

4.3 Cointegration analysis

As integration of order one is established for the common factors of the variables under investigation, the next step is to determine whether cross-member cointegration exists.[6] A long-run relationship between the common components can be investigated using standard time series tests such as the Johansen reduced rank approach (Johansen, 1995). As mentioned before, a small sample size can induce biased realisations of the Johansen test statistics. Hence, this study applies the small sample modification proposed by Reinsel and Ahn (1992) and Reimers (1992), who suggest the multiplication of the Johansen statistics with the scale factor $(T - pk)/T$, where T is the number of observations, p the number of variables and k the lag order of the VAR. This approach corrects for small sample bias such that a proper inference can be made. The empirical realisations of the modified Johansen trace statistic as well as those of the modified Johansen maximum eigenvalue statistic (each

[5]To employ further panel unit root test would certainly be sensible. But as the idiosyncratic components are residuals by definition neither a trend nor a constant is included, restricting the analysis to those tests mentioned above.

[6]Since the panel unit root tests of the idiosyncratic components suggest stationarity, this study do not test for cointegration between the idiosyncratic components.

with and without trend) suggest evidence in favour of a long-run relationship between the common factors of real money, real GDP, the interest rate, the real effective exchange rate and real stock prices (see Table 4). Consequently, cross-member cointegration seems to be actually existent.

Table 4: Results of Johansen's tests for cointegration among common components

H_0	*without trend*				*with trend*			
	Trace Statistic	Critical Value	λ-max Statistic	Critical Value	Trace Statistic	Critical Value	λ-max Statistic	Critical Value
None	78,09*	76,97	45,53*	34,81	77,10*	69,82	44,70*	33,88
At most 1	32,56	54,08	12,44	28,59	32,41	47,86	12,37	27,58
At most 2	20,12	35,19	11,56	22,30	20,04	29,80	11,53	21,13
At most 3	8,56	20,26	5,52	15,89	8,51	15,49	5,51	14,26
At most 4	3,04	9,16	3,04	9,16	2,99	3,84	2,99	3,84

Notes: Potential small sample bias is corrected by multiplying the Johansen statistics with the scale factor $(T - pk)/T$, where T is the number of observations, p the number of variables and k the lag order of the underlying VAR model in levels, see Reinsel and Ahn (1992) and Reimers (1992). Critical values are taken from MacKinnon et al. (1999), and are also valid for the small sample correction. A * indicates the rejection of the null hypothesis of no cointegration at least at the 5% level of significance.

As a next step, this study estimates the established long-run money demand equation for the common components of real money and its determinants. But prior to that, this study investigates the long-run relationship between the original (not decomposed) variables for comparability. In both cases, this study uses the dynamic ordinary least squares (DOLS) estimator proposed by Mark and Sul (2003) who also applied it to panel money demand. The DOLS estimator corrects standard OLS for bias induced by endogeneity and serial correlation. First, the endogenous variable in each equation is regressed on the leads and lags of the first-differenced regressors from all equations to control for potential endogeneities. Then the OLS method is applied using the residuals from the first step regression. The DOLS estimator is preferred to the non-parametric FMOLS estimator because of its better performance. According to Wagner and Hlouskova (2010), the DOLS estimator outperforms all other studied estimators, both single equation estimators and system estimators, even for large samples. Furthermore, Harris and Sollis (2003) suggest that non-parametric approaches such as FMOLS are less robust if the data have significant outliers and also have problems in cases where the residuals have large negative moving average components, which is a fairly common occurrence in macro time series data.

First, the DOLS estimator is applied to the original variables to replicate the established results of the literature on money demand. As a second step, this study presents the estimation of the long-run relationship of the common components of the original variables. The estimated models are:

$$M_{i,t} = \alpha_i + \beta_{1,i} Y_{i,t} + \beta_{2,i} R_{i,t} + \beta_{3,i} EX_{i,t} + \beta_{4,i} W_{i,t} + \varepsilon_{i,t}, \quad \text{and} \tag{9}$$

$$M^c_{i,t} = \alpha_i + \beta_{1,i} Y^c_{i,t} + \beta_{2,i} R^c_{i,t} + \beta_{3,i} EX^c_{i,t} + \beta_{4,i} W^c_{i,t} + \varepsilon^c_{i,t} \tag{10}$$

where $i = 1, ..., N$ refers to each country in the panel and $t = 1, ..., T$ denotes the time period. α_i represents the country-specific fixed effects and the superscript c in equation (10) denotes the common components of the original variables. Since all variables, except the interest rate, are specified in natural logarithms, the estimated long-run coefficients can be interpreted as elasticities and as a semi-elasticity, respectively.

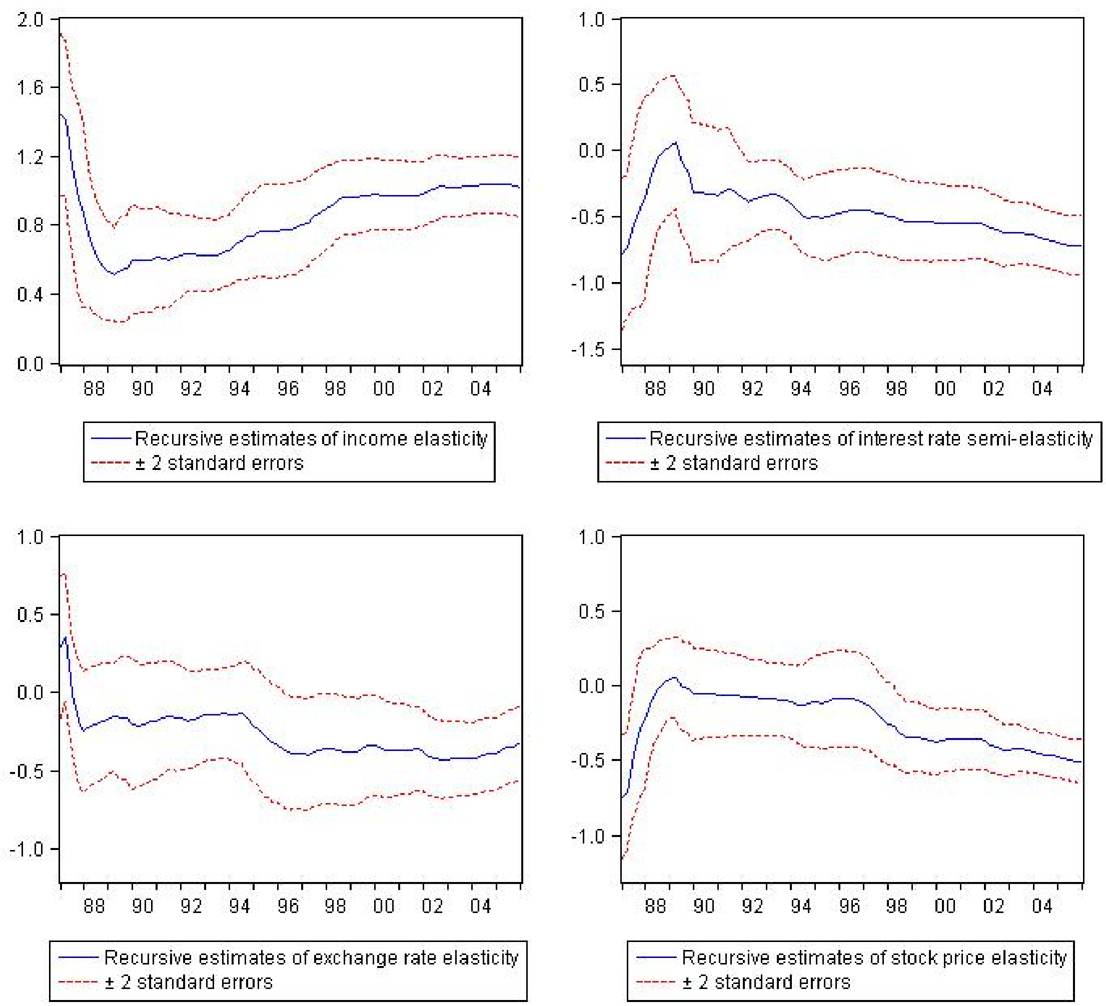
The estimated income elasticity of real money in equation (9) is 1.64 and statistically significant at the 1% level. The finding of an income elasticity greater than one is usual in both time series and panel data studies on money demand. Regarding the interest rate semi-elasticity, this study finds that larger opportunity cost of holding money are connected with lower real balances. More precisely, the short-term interest rate exerts a statistically significant impact on real money of -0.03, where the negative sign is consistent with theoretical postulates. In contrast, the coefficient of the real effective exchange rate has a positive sign (0.14), meaning that a real effective exchange rate appreciation lowers domestic money demand. This result suggests that a possibly negative impact on economic activity exceeds the impact of the income effect. Additionally, the statistical significance of justifies the inclusion of the exchange rate into the money demand function. Furthermore, the statistically significant impact of real stock prices on real money underlines the importance of the inclusion of stock prices in modelling money demand. The corresponding DOLS estimation reports that a 1% increase in real stock prices decreases money demand by 0.15%. Further, the estimated negative impact indicates that the negative substitution effect dominates the positive wealth effect, suggesting that a rise in asset prices reduces the attractiveness of holding money compared to equities. A comparison with the other panel data studies listed in Table 1 reveals that our empirical results for the income elasticity and the interest rate semi-elasticity are actually within the range of previous analyses and show signs that are consistent with money demand. However, the finding that real stock prices are relevant determinants of money demand contradicts those of Arnold and Roelands (2010) and Nautz and Rondorf (2010).

The estimated income elasticity of the common components of real money in equation (10) turns out to be 1.02, close to unity and statistically significant at the 1% level. In fact, a Wald F-test test cannot reject the null hypothesis that the income elasticity is equal to one ($F = 0.40$ [0.53]). Hence, the value predicted by theory can be established by using the common international factors of the variables under investigation. Dreger et al. (2007) propose the very same result as the only further panel data study also supporting long-run income elasticities of money demand using common components. Their estimated income elasticities of the common components are 0.96 and 1.05 whereas their panel income elasticity is clearly greater than one such as in this analysis. Further, this study estimates a coefficient of the common components of real stock prices which is again highly significant and negative. But the coefficient of the common components with a value of -0.50 is absolutely larger than the coefficient of the original stock prices (-0.15). In addition, the interest rate semi-elasticity also rises in absolute terms compared to the previous result (-0.03) and takes a value of -0.71. Again the impact of the interest rate on money demand is negative as anticipated. In contrast to the previous result, the elasticity of real money to the real effective exchange rate is negative in the case of the common components. This time a 1% exchange rate appreciation increases money demand by 0.32%. Thus, the exclusive consideration of the common components without the idiosyncratic component of the underlying variables suggests that the positive currency substitution effect dominates the negative impact on money demand. Moreover, Dreger et al. (2007) also found a positive impact of the exchange rate on money demand. The finding that the coefficients of the common components of the interest rate and stock prices are larger than the coefficients of the original variables may be due to highly integrated financial markets. The established cross-member cointegration already indicates that the stochastic trends are common to all the countries. However, the smaller income elasticity for the common components suggests that the national GDP still plays a major role. Hence, the global business cycle seems not to be as relevant as the international financial integration.

An ADF unit root test verifies the stationarity of the established cross-member cointegration relationship between the common components of real money, real GDP, the interest rate, the real effective exchange rate and real stock prices ($t = -2.98$ [0.04]). Further, the result of the Ramsey RESET test indicates that there are no misspecifications of the long-run equation of the common components

(F = 0.17 [0.68]). At last, stability of the long-run coefficients of the common components of money demand can be detected by recursive coefficient estimates. The stability property of any estimated money demand relation is a critical requirement to ensure the usefulness of such a relation for policy purposes. Figure 1 shows no significant variation in the estimated recursive coefficients as more data is added, suggesting that the long-run money demand coefficients of the common components are stable. In addition, the CUSUM of squares test does not indicate any instability of the residual variance, see Figure 2.

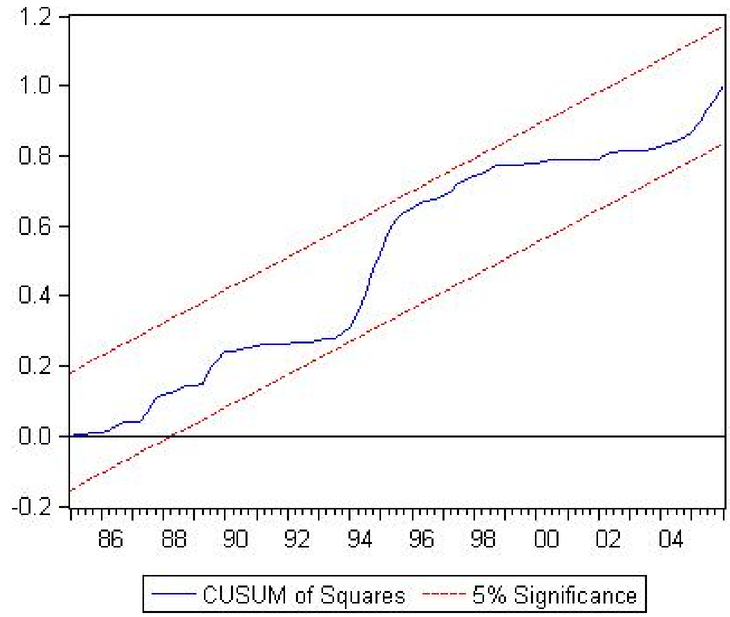
Figure 1: Results of recursive coefficient estimates - Long-run money demand in OECD countries



4.4 Dynamic panel error-correction model

Having established a long-run relationship between the common factors, the next step is to estimate a panel-based error-correction model to determine also the short-run coefficients and the adjustment

Figure 2: CUSUMSQ test - Long-run money demand in OECD countries



coefficient of money demand. A two-step procedure is applied. First, this study employs the long-run equation specified in (10) to obtain the deviation from the established long-run equilibrium of the common components, i.e. $\varepsilon^c_{i,t}$. Then the error-correction model is estimated with the one-period lagged residual from the first step as dynamic error-correction term:

$$\Delta M_{i,t} = \alpha_i + \gamma_{1i}\Delta Y_{i,t} + \gamma_{2i}\Delta R_{i,t} + \gamma_{3i}\Delta EX_{i,t} + \gamma_{4i}\Delta W_{i,t} + \gamma_{5i}\Delta M_{i,t-1} + \lambda_i\varepsilon^c_{i,t-1} + u_{i,t} \tag{11}$$

where Δ denotes the first-difference operator, λ_i represents the speed of adjustment and $u_{i,t}$ is the serially uncorrelated error term with mean zero. Since the sample under investigation includes nearly 100 observations, the usual finite sample bias of dynamic panel estimations, the so-called Nickell-bias (Nickell, 1981), should be negligible. Hence, the use of an instrument estimator such as the GMM estimator proposed by Arellano and Bond (1991) is not required. This study applies the seemingly unrelated regression (SUR) method to incorporate contemporaneous correlation in the errors across equations. SUR estimates the parameters of the system (15) by feasible generalised least squares (FGLS), accounting for heteroskedasticity and correlation between $u_{i,t}$ and $u_{j,t}$, $i \neq j$. It is of particular interest whether national money demand converges to the established common equilibrium path. These long-run dynamics can be studied by testing the significance of the adjustment coefficient, i.e. to check whether the coefficient of the error-correction term represented by λ_i is equal to zero. Table 5 shows the estimated coefficients and the corresponding t-statistics of the panel-based error-correction model.

The estimated coefficients are small but display signs as expected from theory. Changes in income and in the interest rate are estimated to have a highly significant positive (0.323) and negative impact (-0.005) on money demand, respectively. The elasticities of changes in the real effective exchange rate and the real stock prices are insignificant in the short run. Given this result, the currency sub-

Table 5: Short-run coefficients and speed of adjustment of money demand in OECD countries

Independent	SUR system estimated by FGLS		
Variables	Coefficient	t-Statistic	P-Value
Country-specific intercept:			
Australia	0.023	6.283	0.000
Canada	0.019	7.892	0.000
Denmark	0.021	6.663	0.000
France	0.013	5.718	0.000
Germany	0.018	5.991	0.000
Italy	0.020	6.511	0.000
Japan	0.018	8.092	0.000
The Netherlands	0.018	6.951	0.000
Sweden	0.012	2.560	0.011
Switzerland	0.016	3.139	0.002
United States	0.007	3.526	0.000
$\Delta Y_{i,t}$	0.323	4.511	0.000
$\Delta R_{i,t}$	-0.005	-6.368	0.000
$\Delta EX_{i,t}$	0.005	0.613	0.540
$\Delta W_{i,t}$	-0.020	-0.888	0.375
$\Delta M_{i,t-1}$	0.070	2.261	0.024
$\varepsilon_{i,t-1}$	-0.002	-6.186	0.000

stitution hypotheses might hold only in the long run. Furthermore, the estimated coefficient of the error-correction term is highly significant, validating the significance of the cointegration relationship of the common components in the short-run model for money demand. Additionally, the significance of the error-correction term indicates that money demand readjusts towards a common international equilibrium relationship after a shock occurs. The equilibrium is 'international' in the sense that the established long-run relationship is a cross-member cointegration relation and driven by common stochastic trends.

A comparison with the two further panel data studies applying a dynamic error-correction model by Valadkhani (2008) and Nautz and Rondorf (2010) leads to the following conclusions. First, their estimated short-run dynamics are also smaller than the long-run coefficients. Second, Valadkhani (2008) supports that changes in the exchange rate are insignificant in the short run and Nautz and Rondorf (2010) found an insignificant short-run coefficient of stock prices as in this study. Third, the other short-run coefficients estimated in this study are smaller compared to the other studies, except the impact of changes of lagged real money which is within the range of both. The coefficient of the error-correction term might be smaller in this analysis because it measures the speed of adjustment towards an equilibrium relation between common factors but not to an overall equilibrium path. Real money may adjust faster to an equilibrium relation which reflects long-run money demand of not decomposed variables which in addition to the common factors also include the country-specific idiosyncratic components and, thus, promise more explanatory power to changes in real money.

5 Conclusions

This paper studies the long-run money demand function for 11 OECD countries from 1983 to 2006 using panel data and including wealth. The applied factor decomposition provides new empirical insights into the long-run relationship among money and its main determinants. More precisely, the distinction between common factors and idiosyncratic components allows to detect cross-member cointegration and to distinguish between international and national developments as potential drivers of long-run money demand. Indeed, the main empirical finding of this study is that cross-member cointegration is existent and, correspondingly, only the *common* components of real money, real GDP, the interest rate, the real effective exchange rate and real stock prices are cointegrated. This result highlights the relevance of international developments to explain money demand. Hence, policy makers should incorporate cross-country dependencies and international impacts on money demand when designing sensible monetary policy to achieve price stability. Further analysis of the cointegration relationships of (a) the variables under investigation and (b) their common components suggests that the estimated coefficients of the former are within the range of previous panel data studies. The established signs of the income and interest rate (semi-)elasticities are consistent with theoretical postulates in both models. Moreover, the significant negative impact of wealth, represented by stock prices, indicates its importance as determinant of real balances and that the negative substitution effect on money demand dominates the positive wealth effect. However, the relations (a) and (b) differ in the coefficients of the 'financial' variables, the interest rate and the real stock prices, which are larger for the common components. This result highlights that especially the financial markets are highly integrated, since the established cross-member cointegration already indicates a close relation of money and its determinants across the 11 OECD countries. By contrast, the long-run income elasticity of the common components of money demand is smaller than the income elasticity of the original (not decomposed) money demand. Hence, the global business cycle does not seem to be as relevant as the international financial integration. What is more, the stability of the long-run money demand coefficients of the common components can be confirmed by recursive coefficient estimates. This finding is a critical requirement to ensure the usefulness of such a relation for policy purposes. Hence, the long-run money demand relationship between the common components, the cross-member cointegration relation, seems to be a useful reference for monetary policy. Accordingly, the common components of the domestic money stocks may help to reliably identify risks to price stability in addition to the domestic money stocks themselves. Moreover, this paper presents a panel-based error-correction model capturing the short-run coefficients of money demand and, more interestingly, the adjustment coefficient. The estimated short-run coefficients for money demand are smaller than the long-run elasticities and statistically insignificant for the exchange rate and stock prices. Since the residual of the long-run equilibrium relation between the common components is used as dynamic error-correction term, its determined significance means that money adjusts to an international rather than a national equilibrium relationship.

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