Household Money Holdings in the Euro Area: an explorative investigation*

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ABSTRACT

In this paper we analyse household holdings of the broad monetary aggregate M3 in the euro area from 1991 until 2009. Households are the largest money-holding sector in the euro area. We develop four models, two in nominal, two in real terms, with satisfactory economic and statistical properties. The main determinants are a transactions variable, wealth considerations, opportunity costs and uncertainty. In particular housing wealth is found to play an important role. The models are robust to different estimation strategies, samples considered and a multitude of misspecification tests. According to our analysis, it is quite apparent that in equilibrium, households jointly determine consumption and broad money holdings which are both influenced by wealth as well as interest rates. The importance of household money holdings for consumption expenditures may cast doubt on a purely passive role for money.

Keywords: money demand, cointegrated VARs, households

JEL Classification: E41, C32, D12

1. INTRODUCTION

Understanding the demand for money is an important element of a detailed monetary analysis, which aims to extract, in real time, signals from monetary developments that are relevant for the assessment of risks to price stability over the medium to longer term. These longer-term price developments are determined by aggregate money holdings of all sectors. Looking at individual

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money holding sectors, however, may allow the formulation of more consistent and richer explanations of the driving forces for the demand for money as the relative importance of the main motives for holding money - its use as a medium of exchange or as a store of value - varies across sectors. Indeed, heterogeneity in money-holding behaviour goes beyond the sector level to the individual money holder, but harmonised data for a significant sample length is only available at the sectoral level.[2] A sectoral analysis of money demand can, by improving the understanding of the individual components, contribute to a better understanding of the covariation of aggregate money with its determinants.

In general, differences in money demand behaviour may result from two factors:

1. The constraints surrounding the money-holding decision process can vary. This may lead to different elasticities of money demand with respect to the same determinants for individual sectors.
2. The determinants of money demand may differ across sectors, such as alternative investment opportunities and thus different opportunity costs of holding money, or different scale variables.

Consequently, two different modelling strategies need to be considered in the context of sectoral money demand. The first is to estimate money demand using a common set of macroeconomic determinants (see von Landesberger, 2007). This approach allows for a comparison of behaviour across sectors and with aggregate money demand. The alternative modelling approach is oriented toward finding a refined specification for every sector, thus trying to identify the determinants best capable of explaining sectoral money holdings. This is the aim of the present paper which has not yet been done for euro-area data. The understanding of household money holdings is important for several reasons: households are the largest money-holding sector accounting for approximately two-thirds of euro-area M3. They usually hold a large proportion of their money holdings as transactions balances, using these balances mainly as a buffer, while slowly adjusting their portfolio composition. In addition, households' financial decisions are likely to have significant impact on real macroeconomic activity, rendering the interaction between households' money balances and consumption important. The dynamic of household M3 holdings are also found to be informative for price developments in the euro area, giving their explanation a particular relevance for monetary analysis (see European Central Bank, 2006 p. 18).

The paper is structured as follows. The next section provides a review of the literature on household money demand. In part 3, the data and the modelling approach used to estimate the money demand systems are discussed. We also present and discuss the results of four different models of households' M3 demand. Section 4 illustrates the use of these models to understand recent money growth. As a robustness check, section 5 also presents results of a single-equation modelling approach. The last section summarises the findings and provides some implications for monetary analysis.

2. RELATED STUDIES

The following section provides a structured overview of the methods commonly employed in the literature on sectoral money demand and of the main findings. In order to get a better understanding of the results, we distinguish between macroeconomic (time-series) and microeconomic (cross-sectional) studies.

[2] See Martinez-Carrascal and von Landesberger (2010) for a comparison of the behaviour of money demand at the sectoral level and at the micro-economic level for euro-area non-financial corporations.

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2.1. The evidence on household money demand from macroeconomic studies

For the US, the first empirical analysis of the household demand for money was undertaken by Goldfeld (1973). In this study, the demand for M1 is explained by different measures of transactions (GNP and consumption expenditure), controlling for the change in net worth and using the spread between commercial paper and deposit interest rates as opportunity costs. Goldfeld finds that money holdings by households are quite well explained by these variables and have reasonable parameter estimates. Since the publication of Goldfeld (1973), a number of studies have attempted to explain household money demand. In general, these studies have analysed the demand for money by households from a time series perspective using cointegration methods – either based on single equations (Butkiewicz/McConnell, 1995) or based on systems of equations (e.g. Jain and Moon, 1994, Thomas, 1997, Chrystal/Mizen, 2001).

The main scale variable of money demand considered includes real consumer expenditure (Jain and Moon, 1994, Read, 1996, Calza and Zaghini, 2010), measures of real (permanent) disposable income (Butkiewicz and McConnell, 1995, Laumas, 1979), real net labour income (Chrystal and Mizen, 2001) and real GDP (Petursson, 2000, Feiss and MacDonald, 2001). In addition, both real gross personal sector wealth (Thomas, 1997, Read, 1996) and real net total wealth (Chrystal and Mizen, 2001) are intended to capture an additional element of scale.

A variety of interest rate specifications have been tried. These range from simple formulations such as including only the long-term nominal treasury yield (Jain and Moon, 1994), the short-term commercial paper rate (Laumas, 1979), or foreign interest rates (Krupkina and Ponomarenko, 2013). Semi-log and double-log specifications are used (Butkiewicz and McConnell, 1995, Calza and Zaghini, 2010).[3] More complex approaches include the spread between the 3 month t-bill rate and the own rate of money (Thomas, 1997, Petursson, 2000) or between the yield on public bonds and the own rate (Read, 1996). Chrystal and Mizen (2001) even include two interest rate terms in their model, the rate on savings deposits minus a money-market rate and the spread between the rate on consumer credit and the base rate. An additional variable repeatedly included in models for the UK is the rate of inflation, reflecting either the return on real alternatives to money or helping to test for price homogeneity (Thomas, 1997, Chrystal and Mizen, 2001, Feiss and MacDonald, 2001).

The main findings are that household real balances are cointegrated with measures of income and interest rates. Several studies emphasise, both for narrow and for broad monetary aggregates, a transactions-based explanation of money demand (Jain and Moon, 1994), captured by a strong interaction between household money holdings and consumption (Thomas, 1997). Broadening the analytical framework to include households demand for loans, Chrystal and Mizen (2001) find that consumption, money holdings and credit interact both in the determination of the long-run equilibrium and in their short-run adjustment. Read (1996) provides evidence for Germany that households' money holdings tend to be determined by longer-term considerations, whereas the corporate sector is far more responsive to short-term influences.

2.2. Microeconomic evidence at the household level

While the focus of this paper is a time series perspective, evidence brought forward in cross-sectional studies could potentially contribute valuable further insights in the specification of the models. The monetary data examined in these studies is generally taken from household surveys. A first study was conducted by Garver and Radecki (1987) on a cross-sectional sample of US data. They investigate the holdings by households of a narrow measure of money consisting of currency holdings plus total checking accounts. The scale variable considered is total household

[3] Calza and Zaghini (2010) find that the welfare costs of inflation differ for households and firms and crucially depend on the double-log versus semi-log specification.

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annual income, while the opportunity costs of holding money are measured by the average money-market deposit rate minus the rate of interest earned on checking accounts. A number of dummy variables are included to account for different types of checking accounts. The study emphasises the transactions motive for holding money and support the use of the macroeconomic approach to the demand for narrow money. Attanasio et al. (1998) also investigate households' holdings of real cash balances using non-durable consumption as scale variable and an interest rate as opportunity costs. The interest rate and expenditure elasticities found for the demand for cash are close to the theoretical values implied by standard inventory models. With data for Japan, Fujiki and Hsiao (2008) constitute an exception by examining the issues of unobserved heterogeneity among cross-sectional units and stability of an aggregate function for broad money. The estimated income elasticity for Japanese household M3 is around 0.68 and the five year bond interest rate elasticity is about -0.12. Anderson and Collins (1997) investigate M2 growth in the United States from 1990 to 1993 using a model of household demand for liquid wealth. The authors find that the own-price elasticity of money demand rose substantially during this period and report sizeable cross-price elasticities of money with respect to other liquid financial assets, notably with mutual funds. They also suggest that households may respond more rapidly to changes in market interest rates than is often assumed. Tin (2008) examines the precautionary demand for transactions balances. The monetary measure considered is non-interest-earning checking accounts by households in the US. The study indicates that income volatility is a significant determinant of money holdings as predicted by the inventory theory of money demand. The relative magnitudes of the elasticities of income and income volatility suggest that the strength of the relationship between the precautionary motive and money demand is much weaker than the strength of the relationship between the transactions motive and money demand. Unfortunately, cross-sectional studies have not yet investigated holdings of components of broad monetary aggregates.

3. THE EMPIRICAL APPROACH

In what follows, we try to model euro area M3 holdings by households using a broad set of explanatory variables. In line with the literature reported above, the aim is to estimate a money demand relationship using system cointegration techniques.

3.1. The framework and the data

Monetary theory suggests different determinants for the holding of broad money, which like for other financial assets, is part of a portfolio allocation decision (see Friedman, 1956, Tobin, 1969). At least some of the assets included in broad money in addition provide liquidity services to their holder. A general formulation of the determinants can be stated:

$$m = \beta_1 p + \beta_2 y + \beta_3 w + \beta_4 i^{alt} + \beta_5 i^{own} + \beta_6 \sigma$$

whereby m denotes the stock of money, p the price level, y and w the level of transactions and wealth, respectively, i^{alt} and i^{own} the returns for investments outside M3 and in monetary assets included in M3 and σ represents variables capturing different aspects of uncertainty, be they economic, financial or geopolitical. The β_i's ($i = 1,\ldots,6$) denote the parameters capturing the effect of the respective determinants on money holding.

Three key economic features have to be fulfilled by empirical estimates in order to identify a money demand function:

1. β_2 and β_3 must be positive,
2. β_4 must be negative and β_5 positive,

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3. discrepancies between actual money and equilibrium holdings lead to an adjustment in money growth.

For the purpose of our analysis, *m* is broad money holdings of households. The sector also comprises non-profit institutions serving households.[4] M3 data is taken from the official ECB database for the period since 1999.[5] The criticism is sometimes made that sectoral money demand studies for the US, which are based on flow-of-funds data, might be affected by the fact the household money holdings are a residual position in the data. In the euro area, over the sample considered, this is not the case, as between 81% and 88% of M3 data, namely all deposits (including repurchase agreements) held by the household sector were directly reported by MFIs.

Two measures of the price level are considered potentially relevant for households to calculate real balances: the private consumption deflator and the harmonised index of consumer prices. The scale of households' transactions settled using money may be captured by a variety of variables. Following the literature, the variables considered are the level of consumption expenditures, disposable income, a measure of household expenditures consisting of consumption plus investment into housing as well as measures of household wealth. Boone et al. (2004), Greiber and Setzer (2007), Beyer (2009) and de Bondt (2009) find a significant role for wealth in euro area money demand, with the latter three studies emphasising the role of housing wealth. A number of wealth measures are therefore considered, specifically gross total household wealth and its components gross financial household wealth and gross housing wealth, as well as net total household wealth. In addition, a measure of longer-term housing wealth using the trend in house prices is used. The reason underlying such a calculation is that households may not perceive themselves to become more or less wealthy with high frequency movements in the prices of their asset holdings, but rather take a medium term-view of asset prices.[6]

In order to model the opportunity costs of holding money, a wide range of alternative returns and interest rates are initially selected: these include the long-term interest rate on bank lending to households for house purchase, the yield on long-term government bonds, the yield on corporate bonds and a short-term money market interest rate. While the last three interest rates can be seen as fairly common choices, the consideration of the bank lending rate as an alternative investment opportunity for households rests on the observation that in the presence of intermediation costs between borrowing and lending from a bank, a reduction in borrowing generally offers households a better return than holding money. Thus, the use of a bank lending rate draws on the notion that the household sector holds money as a buffer stock which will be reduced as the financing cost of households increases.

What is left are proxies for risky investments: the dividend yield as well as the earnings yield of euro area non-financial corporations are considered to take developments on stock markets into account. Friedman (1988) outlines the interactions between money holdings and the stock market. In addition to the realised earnings per share, following the approach proposed by Chordia and Shivakumar (2002) as well as Stern and Stern (2008), expected earnings per share are estimated based on a regression relating the earnings on equity on the recent dividend yield, the real short-term interest rate, the slope of the yield curve and the spread between corporate and government bonds. Moreover, the simple price/earnings ratio is also included in the data set. It is determined by expectations about discount rates and about earnings growth, with the former mainly influencing the evolution in the long-run (see Fama and French, 2002, Campbell and Shiller, 1998). It can therefore be considered as a proxy for the discount rate applied to investments in risky assets. The spread between corporate and government bonds can also be viewed as a proxy for risk. The

[4] The level of money stock is the notional stock adjusted for seasonal effects with Tramo-Seats. The data is extended backwards before 1999 Q1 assuming an unchanged sectoral share in money market funds, currency in circulation and debt securities holdings at the levels of 1999 Q1. These instruments represent only a small share of household M3 holdings in 1999.

[5] The overall approach to the construction of the series is described in ECB (2006b).

[6] The trend in house prices is derived using an approach common to the analysis of the link between money and asset prices (Detken and Smets, 2004, Adalid and Detken, 2007). The trend is estimated using a very slow adjusting HP-Filter ($\lambda = 100{,}000$).

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return on monetary assets is captured by the own rate of households' M3 holdings, calculated as a weighted average of the remuneration of the instruments included in M3.

Interest rates can enter the money demand relationship in two functional forms: first, the semi-log specification, which is the most popular in money demand studies (see, e. g., Ericsson, 1998). It estimates semi-elasticities and implies the same response of money holdings to each percentage point reduction in nominal interest rates. Second, the double log form proposed, inter alia, by Lucas (2000). It entails that a percentage point reduction in nominal interest rates has a proportionally greater impact upon money holdings the lower the level of interest rates, i.e., the semi-elasticities vary with the level of interest rates. For higher levels of interest rates the two functional forms lead to similar results. The non-linear impact at low levels of interest rates can be motivated by the prevalence of fixed costs into alternative investment opportunities and that households who hold only cash do not incur this cost. A logarithmic money demand function may also be rationalized within a stylised general equilibrium model with money (Chadha et al., 1998, Stracca, 2001).

Finally, measures of uncertainty that proxy households' economic and financial confidence are also included. The measures considered are the EU Commission's index of consumer confidence and its subcomponents (e.g. employment expectations) as well as the actual rate of unemployment, which was found relevant by de Bondt (2009). Furthermore, financial market uncertainty in the form of stock and bond market volatilities (Carstensen, 2006) and of the uncertainty factors estimated by Greiber and Lemke (2005) is taken into account. The latter derive composite series for uncertainty using an unobserved components model. One of their indicator variables is mainly based on financial market data, such as medium-term returns, loss and volatility measures while the other factor is more heavily geared toward business and consumer sentiment. Both the individual economic variables as well as the aggregate factors are intended to capture the economic forces impacting on the household's decision to hold money for precautionary reasons.

The set of explanatory variables presented above and shown in Chart A in annex 1 allow the specification of a whole battery of equations. A number of alternative specifications for household M3 holdings are tested and a selection of the most promising specifications is presented in more detail below. The equations are chosen to get economically plausible specifications and statistically sound estimation results. More specifically, the equations considered are:

Model 1-n: $$m = f\left(\overset{+}{pc},\ \overset{+}{rc},\ \overset{+}{rthw},\ blr \overset{-}{-} own,\ \overset{+}{GL1},\ \overset{?}{UN^e}\right)$$

Model 2-n: $$m = f\left(\overset{+}{pc},\ \overset{+}{rc},\ \overset{+}{rtw},\ blr \overset{-}{-} own,\ p \overset{-}{-} e\right)$$

Model 3-r: $$m - pc = f\left(\overset{+}{rdi},\ \overset{-}{dpc},\ IRL \overset{-}{-} IRS,\ p \overset{-}{-} e\right)$$

Model 4-r: $$m - pc = f\left(\overset{+}{rc},\ \overset{-}{dpc},\ IRL \overset{-}{-} OWN,\ p \overset{-}{-} e,\ \overset{?}{C^e}\right)$$

where variables written in lower case letters enter the VAR systems in logarithms. The sign above the variables indicates the theoretical expected impact.

Model 1-n explains *nominal* household M3 holdings using the private consumption deflator *pc* and two scale variables, real private consumption *rc* and a measure of the trend in housing wealth deflated with the private consumption deflator *rthw*. The spread between the bank lending rate for house purchases *blr* and the own rate on households' M3 holdings *own* (both in logs) enters the money demand model as the measure of opportunity costs. In order to model precautionary motives of the demand for money, the uncertainty measure developed in Greiber and Lemke *GL1*

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related to capital market forces enters the model as a measure of uncertainty. Finally, expectations with regard to unemployment over the coming twelve months UN^e from the survey of the EU Commission are included in the VAR system. A deteriorating employment situation may, on the one hand, induce households to hold greater money balances to meet unforeseen expenditures. On the other hand, the expected deteriorating economic environment and increasing uncertainty may reduce the attractiveness of nominal assets and induce the purchase of more real assets. Therefore, the total impact on the demand for money is ambiguous (see Atta-Mensah, 2004b).

Model 2-n draws on a similar set of variables as Model 1-n, but includes total household wealth deflated with the private consumption deflator as the relevant wealth measure *rtw*. Precautionary money holdings are captured by the price earnings ratio on euro area equity $p - e$ which may be considered as a measure of risk on financial markets.

Model 3-r explains *real* household M3 balances with only one scale variable – real disposable income *rdi*, but includes two measures of opportunity costs, the change in the consumption expenditure deflator *dpc* and the term spread $IRL - IRS$. As in Models 2-n, the price-earnings ratio $p - e$ is also included.

Model 4-r builds on the previous model, but substitutes *rdi* with real consumption expenditure *rc*, $IRL - IRS$ with the spread between the long-term nominal bond yield and the own rate of household M3 holdings $IRL - OWN$. In addition, expectations of economic prospects and thus future consumption are taken into account. In order to capture this forward-looking element, expectations with regard to the strength of economic activity from the EU Commission consumer confidence surveys C^e are included.

3.2. Overview of the modeling outcomes

The empirical analysis is conducted on seasonally adjusted quarterly data over the sample period 1991 Q1 to 2009 Q3.[7] The estimations are performed over the shorter sample 1991 Q1 to 2008 Q3 in order to avoid any contamination of the results from the financial market crisis following the default of Lehman Brothers in September 2008, with the last four observations analysed in Section 4.

To determine the order of integration of the time series, ADF and KPSS tests are carried out (see Table 9 in annex 1). The two tests – together with conceptual considerations for some of the borderline cases on the boundedness of the variance - support the view that most series in levels, except the spreads, are I(1). An additional test for stationarity of the variables within the cointegrated VAR supports this decision (see Table 11 in annex 1). That said, it should be recognised that some variables may still exhibit quite persistent fluctuations in first differences. Difference stationarity of money and prices may be considered slightly at odds with a part of the recent empirical literature on money demand that finds these variables to be I(2).[8] Given its prominence in the empirical money-demand literature, this possibility is entertained in the modelling approach applied below.

Within our cointegrated VAR approach the first step consists in estimating an unrestricted VAR system comprising an endogenous variables vector y_t and exogenous (non-modelled) I(0) variables vector x_t:

[7] Davidson and MacKinnon (1993, p. 714) prove that unit root test statistics are biased against rejecting the null hypothesis when working with seasonally adjusted data. As nearly all our variables are clearly *I(1)* (see annex 1) this reduces the severity of this problem. Furthermore, Ericsson et al. (1994) show theoretically and empirically within the Johansen framework that the number of cointegrating vectors and the cointegrating vectors themselves are invariant to the use of seasonally adjusted or unadjusted data.

[8] See Juselius (2006) and for instance Feiss and MacDonald (2001).

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$$y_t = \sum_{i-1}^{p} \Pi_i y_{t-i} + \Psi_0 x_t + \Phi D_t + \varepsilon_t \qquad (1)$$

The errors ε_t are assumed to be $NI\sim(0,\Omega)$. Π_i and F are matrices containing the parameters of the model. D_t is a vector of deterministic variables, potentially comprising constant terms μ_0 or deterministic trends. Given the quarterly data used, the maximum lag length p is set equal to four in order to determine the appropriate number of lags for each model. The Akaike information criterion (AIC) is used to select the lag length for conducting the remainder of the analysis and the outcome is cross-checked with Likelihood Ratio tests (see Table 10 in annex 1). The AIC tends to favour the inclusion of more lagged terms than for example the Schwartz information criterion.[9] Overestimation of the order of the VAR is much less serious than underestimating it, as shown for example by Kilian (2001). In the models presented below, the lag length retained ranges between two and three in levels.

Table 1 presents the outcome of standard specification tests of the respective VAR systems. The null of no autocorrelation in the residuals cannot be rejected in any of the systems at conventional significance levels. In a similar vein, tests for ARCH effects in the residuals are also not significant. By contrast, non-normality of the residuals is detected for two models owing to the presence of outliers.[10]

Table 1
Residual properties for the VAR systems

	Test statistic	p-value		Test statistic	p-value
Model 1-n					
LM-AR(1)	F(36,54) = 0.69	0.95	Multivariate ARCH	F(441,61) = 1.11	0.32
LM-AR(4)	F(36,51) = 0.99	0.50	Normality	F(12,59) = 1.67	0.14
Model 2-n					
LM-AR(1)	F(36,48) = 1.16	0.31	Multivariate ARCH	F(441,55) = 1.13	0.31
LM-AR(4)	F(36,44) = 0.59	0.95	Normality	F(12,53) = 2.27	0.02
Model 3-r					
LM-AR(1)	F(25,57) = 1.53	0.09	Multivariate ARCH	F(225,50) = 1.32	0.09
LM-AR(4)	F(25,54) = 0.88	0.63	Normality	F(10,61) = 0.79	0.64
Model 4-r					
LM-AR(1)	F(36,53) = 0.96	0.55	Multivariate ARCH	F(441,60) = 1.07	0.39
LM-AR(4)	F(36,50) = 0.65	0.91	Normality	F(12,58) = 2.58	0.01

Note: p-values derived from comparison with respective asymptotic distribution.

In a second step, we reformulate the VAR system into a VECM and test for the rank of the matrix Π_1 using the trace test (see Johansen, 1996):

$$\Delta y_t = \Pi_1 y_{t-1} + \mu_0 + \sum_{i=1}^{l-1} \Gamma_i \Delta y_{t-i} + \Psi_0 x_t + \varepsilon_t \qquad (2)$$

where l indicates the lag length determined in the previous step. The trace tests were conducted assuming the presence of a linear deterministic trend in the time series and a non-zero intercept

[9] Lütkepohl and Saikonnen (1997, p. 16) find that "In most cases AIC and HQ have a slight advantage over the very parsimonious SC criterion".

[10] While normality of residuals is part of the theoretical assumptions of the distribution of residuals, the violation of normality may not be a severe deficiency as the evaluation of the trace test will be supported by bootstrapping results.

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m_0 in the cointegration relationship.[11] Table 2 reports the trace test statistics for different rank assumptions as well as the p-values obtained from comparing this test statistic with the critical values derived by MacKinnon et al. (1999).[12] All models reject the rank 0 at the 5% significance level, with model 1 also rejecting rank 1 and 2. However, given the presence of exogenous I(0) regressors in one of the models (in model 1-n unemployment expectations) and the small sample size, caution in assessing the number of long-run relationships possibly present in the data using this metric seems reasonable. Therefore, more informative parametrically bootstrapped p-values generated from 1,000 replications are undertaken.[13] While the theory on bootstrapping in a non-stationary framework, such as the cointegrated VAR, is still undiscovered territory, the usual theoretical properties from models with stationary variables seem to apply in this setting as well (Juselius, 2006, p. 157, Swensen, 2006).[14] At the 10% significance level, all model specifications indicate one cointegration relationship.

System cointegration tests are well-known to have low power. This gives reason to believe that such tests have a tendency to favour the choice of too few long-run relations. Juselius (2006) suggests the use of as much additional information as possible in the rank determination. We follow this lead and additionally:

1. examine whether the t-value on the load factor of an additional cointegration vector is less than 2.6;
2. analyse recursively the trace statistic and the cointegration relations;
3. check the economic interpretability of the results.
4. While the first and third approaches require the specification of the cointegrated VAR systems, the second approach can be generated on the basis of the unrestricted VAR model.

Table 2
Trace test results

Model		Rank					
		0	1	2	3	4	5
1-n	test statistic	133.56	83.98	46.95	13.22	5.71	0.83
	p-value	0.000	0.001	0.020	0.671	0.475	0.362
	bootstrapped p-value	0.022	0.213	0.473	0.961	0.792	0.658
2-n	test statistic	119.16	53.99	31.21	15.61	7.11	1.90
	p-value*	0.000	0.462	0.655	0.738	0.565	0.17
	bootstrapped p-value	0.074	0.798	0.785	0.957	0.854	0.292
3-r	Test	75.77	37.75	14.07	6. 52	1.863	-
	p-value*	0.016	0.313	0.837	0.633	0.172	-
	bootstrapped p-value	0.012	0.263	0.789	0.712	0.215	-
4-r	Test	98.93	58.13	31.71	15.41	4.42	0.32
	p-value*	0.029	0.297	0.628	0.752	0.866	0.573
	bootstrapped p-value	0.052	0.336	0.601	0.768	0.867	0.536

* Barlett corrected trace statistic.

[11] The cointegration analysis and the results presented in the remainder of this note are computed with the Structural VAR software which was kindly provided by Anders Warne. See http://www.texlips.net/svar/source.html.

[12] Where no exogenous I(0) regressors are included in the VAR systems, the Bartlett correction of the test statistic is undertaken and compared with the critical value.

[13] The parametric bootstrapping procedure implies drawing new innovations from a multivariate standard normal distribution. These innovations are then transformed into bootstrapped residuals using the estimated covariance matrix from the original estimated residuals. On the basis of the initial values and taking the estimated parameters as given, new data series are constructed and the model re-estimated on the new data set. An alternative would be to adjust the test statistics (see, e. g., Reimers, 1991) or the critical values (see Cheung and Lai, 1993).

[14] In particular, a bootstrapped statistic can be expected to have errors in null rejection probabilities that are of a smaller order of magnitude, as the sample size goes to infinity, than its asymptotic analogue when the asymptotic distribution of the statistic is invariant to the parameters of the model. Almost all statistics that we bootstrap are invariant in this sense. See Park (2005) and Chang et al. (2002) for some recent developments regarding models with unit roots.

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In order to robustify the modelling decision on the basis of the trace test, Chart B in annex 1 shows the recursively estimated trace test statistic for the hypothesis of rank one. For the nominal models, the outcome is more reassuring than for the models specified in real terms. But in any case, the trace statistic is fairly stable and around the level of the critical value for the 5% confidence level. Therefore, in the following, a rank of one is assumed for modelling the VAR systems. This decision is also supported by Chart C in annex 1, which presents recursive estimates of the largest eigenvalue for a given set of parameters of the short-run and deterministic variables. For all four models, the depicted eigenvalue bands do not cross the zero line.

Parameter stability has been an issue of primary concern in the context of money demand estimations. Table 3 presents the outcome of tests on parameter non-constancy under the retained assumption that the P-matrix has rank 1. The Ploberger-Krämer-Kontrus (1989, henceforth PKK) fluctuation test examines the constancy of the parameters capturing the short-run dynamics. The test is conducted for all individual equations of the VAR system, but the table reports only the outcome of the money demand equation. The PKK test is unable to reject the null of parameter constancy which supports the eigenvalue analysis reported above. In addition, Table 3 also shows the results of the Nyblom tests for possible non-constancy of the parameters of the cointegration vector. Again, the stability of the parameters is not rejected.

Table 3
Stability tests – Nyblom & PKK (Rank 1)

Model 1-n	Test statistic	p-value
Nyblom Sup F	87.80	0.16
Nyblom Mean	21.51	0.13
PKK S(9)	0.66	0.96
Model 2-n		
Nyblom Sup F	24.99	0.25
Nyblom Mean	6.06	0.31
PKK S(14)	1.72	0.28
Model 3-r		
Nyblom Sup F	3.75	0.31
Nyblom Mean	0.73	0.58
PKK S(7)	0.56	0.98
Model 4-r		
Nyblom Sup F	28.39	0.54
Nyblom Mean	8.52	0.54
PKK S(8)	0.86	0.76

Note: p-values derived from comparison with respective bootstrapped distributions. PKK is calculated on equation-by-equation basis.

Moreover, tests on the stationarity of the variables included in the VAR are conducted to determine whether the reduced rank of the P-matrix resulted from the inclusion of stationary variables. Table 11 in annex 1 reveals that the reduced rank does not seem to result from a single stationary variable. This again supports our decision to treat the variables as I(1).

In order to get further insights into the relationship between the variables and to help the identification of the cointegrated VAR system, we run joint weak exogeneity tests on the variable set. The tests also help to detect the common driving forces amongst the variables of the system.

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A weakly exogenous variable contributes to the common trend of the other variables in the VAR system. At the same time, shocks to variables that are not weakly exogenous have no permanent effect on any other variable in the system. Table 4 shows the outcome of this analysis. In all four models, the assumption that the cointegration relationship does not affect household M3 balances is clearly rejected (see for further discussion the presentation of the models below).

Following the choice for the rank of the Π_1-matrix in (2), finally α cointegrated VAR system is estimated. This entails the identification and estimation of the vector of load factors a and the cointegration vector β' in (3)

$$\Delta y_t = \alpha\beta' y_{t-1} + \mu_0 + \sum_{i=1}^{l-1} \Gamma_i \Delta y_{t-i} + \Psi_0 x_t + \varepsilon_t \tag{3}$$

Table 4
Tests for weak exogeneity of variables

	Null hypothesis: α in equation k is zero Alternative hypothesis: α in equation k is not zero							
	Model 1-n		**Model 2-n**		**Model 3-r**		**Model 4-r**	
Equation for	F(1,58)	p-value	F(1,52)	p-value	F(1,60)	p-value	F(1,59)	p-value
m	26.85	0.00	13.35	0.00	-	-	-	-
$m - pc$	-	-	-	-	5.16	0.03	15.33	0.00
pc	0.60	0.44	13.03	0.00	-	-	-	-
Rc	10.49	0.00	16.72	0.00	-	-	0.17	0.68
Rdi	-	-	-	-	20.25	0.00	-	-
Rtw	-	-	5.19	0.027	-	-	-	-
$Rthw$	1.81	0.18	-	-	-	-	-	-
Dpc	-	-	-	-	5.04	0.03	0.17	0.68
$blr - own$	0.41	0.52	7.66	0.01	-	-	-	-
$IRL - IRS$	-	-	-	-	1.51	0.22	-	-
$IRL - OWN$	-	-	-	-	-	-	1.39	0.24
$GL1$	0.42	0.52	-	-	-	-	-	-
$p - e$	-	-	0.12	0.73	4.15	0.05	2.43	0.12
C^e	-	-	-	-	-	-	16.64	0.00

The results for the β' and α vectors are presented in the next sections. The long-run relationships are checked for robustness in Section 5.

3.2.1. The long-run relationships – the bs

Models 1-n and 2-n describe nominal M3 balances of households, while models 3-r and 4-r determine real balances. The models differ in terms of explanatory variables. Table 5 shows the point estimates of the parameters.

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Table 5
The restricted cointegration vectors *β*

Model 1-n								Test
	m_{t-1}	pc_{t-1}	rc_{t-1}	$rthw_{t-1}$	$(blr - own)_{t-1}$		$GL1_{t-1}$	F(2,59)
	1.000	-1.000	-0.67	-0.67	0.70		-1.23	0.65
			[0.04]	[0.04]	[0.09]		[0.28]	[0.53]
Model 2-n	m_{t-1}	pc_{t-1}	rc_{t-1}	rtw_{t-1}	$(blr - own)_{t-1}$		$(p - e)_{t-1}$	F(3,54)
	1.000	-1.000	-0.5	-0.5	0.08		-0.23	1.23
			[0.04]	[0.04]	[0.02]		[0.02]	[0.31]
Model 3-r	$(m - pc)_{t-1}$		rdi_{t-1}		$(IRL - IRS)_{t-1}$	dpc_{t-1}	$(p - e)_{t-1}$	-
	1.000		-1.82		0.07	0.23	0.51	-
			[0.016]		[0.02]	[0.08]	[0.08]	
Model 4-r	$(m - pc)_{t-1}$		rc_{t-1}	C^{e}_{t-1}	$(IRL - OWN)_{t-1}$	dpc_{t-1}	$(p - e)_{t-1}$	F(2,60)
	1.000		-1.000	-0.007	0.08	-	0.46	2.61
				[0.001]	[0.02]		[0.08]	[0.08]

Note: *α* restricted as in Table 6, standard errors in square brackets.

Theoretically, money holdings should be linear homogenous in the price level in the long-run, thus suggesting to impose a parameter restriction of -1 on the long-run parameter for the price level. At the same time, the consumption expenditure deflator used in the empirical analysis might be a restrictive proxy for the price level actually entering households' money holding decisions. In this case parameter estimates larger than one might also be justified. For both nominal models, the assumption of linear homogeneity is not rejected by the data.[15] Linear homogeneity between household balances and prices permit the reformulation of the models in terms of a demand for real money balances. The results of such a reformulation are presented in Annex II (model 1-r), with most features of model 1-n remaining unchanged.[16] This finding provides an empirical justification to impose the restriction from the outset and estimate models 3-r and 4-r in terms of real balances.

Turning to the ***parameter estimates on the scale variables***, i.e. consumption, income and wealth, the following restrictions are proposed:

1. In the case of model 1-n, an over-identifying restriction is introduced by postulating that real consumption and trend housing wealth are equally important for the demand for money, an assumption similar to Thomas (1997). This equality restriction is not rejected. Together, the parameters sum to 1.32, a value not out of line with results from analyses with euro area aggregate M3 (see for instance, Calza et al., 2001). A restriction to the value of unity is however rejected.
2. However, a slightly different constraint implying that the ratio of broad money to consumption is determined by the ratio of total wealth to consumption can be imposed in model 2-n. Indeed, the parameters of real consumption and real total wealth can be constrained to an equal weighting of 0.5 on each variable and thus sum to unity.

[15] The test statistic is distributed as F(1,59) = 0.31 (p-value = 0.58).

[16] Indeed, except for the change in prices which enters the long-run relationship with a negative sign, in-line with an interpretation as an opportunity cost, most features are very similar. This suggests that the deviation from linear homogeneity in model 1 can be considered of second order.

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3. In model 3-r, no over-identifying restrictions are imposed on the cointegration relationship. The parameter estimate obtained for real disposable income of +1.8 is very high, but the model includes only disposable income as a scale variable. A parameter estimate in this order of magnitude is generally thought to reflect the fact that households hold broad money above and beyond transaction purposes, for instance, as a store of wealth.
4. Model 4-r takes a different approach to explain the scale of real household M3 balances. Instead of disposable income, it focuses, like the nominal models, on consumption expenditure. Additionally, expectations of future economic activity (C^e) enter the system. The other variables are similar to model 3-r. A major difference between the two real models is the parameter estimate for real consumption and real disposable income. While the parameter for consumption in model 4-r can be constrained to unity, this is strongly rejected for the parameter of disposable income in model 3-r. Therefore, this parameter restriction seems due to the inclusion of expectations on future economic activity, which substitute for some of the explanatory power of real disposable income in model 3-r.

The parameter estimates of the ***opportunity cost*** variables exhibit the following characteristics:

1. In models 1-n and 2-n, the opportunity costs of holding money are proxied by the spread between a bank lending rate and the own rate of household M3 balances (in logs). In both models, it has the expected sign and is significantly different from zero (at the 5% percent level). However, the parameter estimate in model 2-n is much smaller than in model 1-n (by a factor of 10).
2. In model 3-r, two proxies for opportunity costs of holding money are included: first, the slope of the yield curve is found to have a negative impact on the level of real M3 holdings. Second, consumer price inflation d*pc* has a negative impact on real household money holdings. This is also present in Thomas (1997) and in Coenen and Vega (1999). Restricting the inflation rate parameter to zero is rejected by the data.[17] This finding suggests that inflation is a relevant opportunity cost for households in the long-run, as households shift out of money and into real assets with a higher level of inflation. Comparing the parameters of the slope of the yield and inflation suggests that the substitution between money and real assets may be significantly stronger than between financial assets.
3. By contrast, in model 4-r the parameter on inflation can be constrained to zero, while the parameter estimate for the spread between the long-term bond yield and the own rate on household M3 holdings has a similar magnitude as in model 3-r.

Turning to the variables intended to capture ***precautionary considerations***, the following observations can be made:

1. The sign on the financial market uncertainty measure in the long-run relationship has the expected positive sign and is significantly different from zero, implying that higher financial market uncertainty leads to higher money holdings.
2. In addition, model 1-n also has as an exogenous regressor in the system's short-run dynamics, households' unemployment expectations in the coming twelve months as a proxy for consumer confidence. The point estimate is negative which is in line with Atta-Mensah (2004b) for Canada. Obviously, this reflects the fact that over the sample the effect via precautionary money holdings dominates.
3. In models 2-n, 3-r and 4-r, the price earnings ratio exerts a negative effect on household money holdings, in line with an interpretation that emphasizes the implicit discount factor embodied in it. When corporate earnings relative to observed stock prices are high, money holdings are high owing to the high uncertainty as reflected in the implicit strong discounting of earnings (and vice versa).

[17] $F(1,61) = 11.21$, p-value = 0.00.

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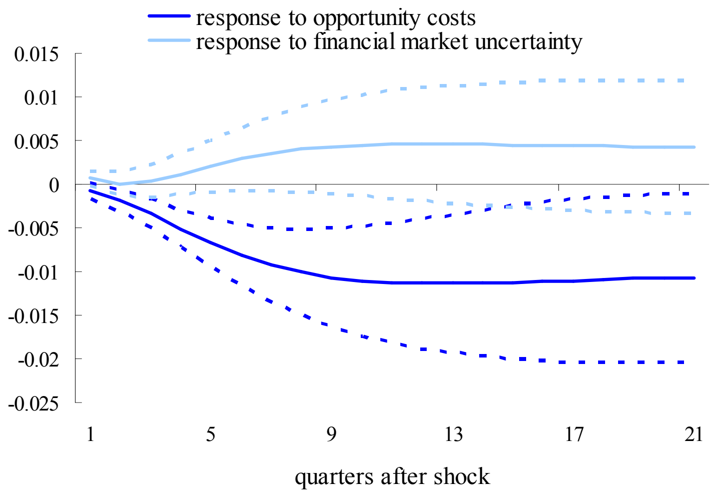
In order to highlight the countervailing impact from portfolio considerations on household money holdings, Figure 1 shows the generalized impulse responses of household M3 to one standard deviation shock in the opportunity cost and uncertainty variables in the context of Model 1-n as an example. A widening of the spread has a significant and negative effect on the level of real M3 holdings by households. In the face of higher borrowing costs, households have an incentive to reduce their holdings of the lower-yielding monetary assets. The complete impact has unfolded after around 10 quarters and remains negative thereafter on the level of money. An increase in the level of financial market uncertainty implies higher money holdings, with the effect taking around 10 quarters to unfold as well. In terms of magnitude, the impact of the interest rate seems to dominate uncertainty effects in this model.

Figure 1
Generalized impulse response of household money to opportunity costs and uncertainty



Note: Dotted lines denote 95% confidence interval around the respective impulse response.

3.2.2. The cointegration relationships – the $\beta' y_t$

The cointegration relationships of all four models are shown in Chart F in annex 1. The charts illustrate quite persistent deviations from the embodied "equilibrium" level. In the case of model 1-n, downside deviations from the average level are observed for periods when the pace of economic activity was slowing (1992–1994, 2001–2003 and since 2007 Q1), while upside deviations are observed particularly for the period 1995–1996 and 1999–2000 before the burst of the dotcom bubble and to a lesser extent between 2004 and 2006.

The co-integration relation from model 2-n exhibits a visibly different pattern from that obtained from model 1-n, especially for the most recent period between 2004 and 2008. The cointegration relationship suggests that M3 holdings have been broadly in line with the level implied by the longer-term determinants for this period and does not point to the sharp decline visible in model 1-n since the end of 2006. A casual inspection of the co-integration relation of model 2-n suggests a break in the series around year 2000. However, checking for parameter stability of the long-run relationship with the Nyblom test and the one-step ahead forecast Chow test does not suggest instability in the parameters of the M3 equation, even in 2007/08 (see Figure 2a). Occasional predictive failures may not be a reason for concern, as these may arise when major shocks occur to the system, while the prediction tests might be useful as a diagnostic tool for parameter stability over a longer time period (Juselius, 2006, p. 164).

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Figure 2a
Chow test for household M3 equation

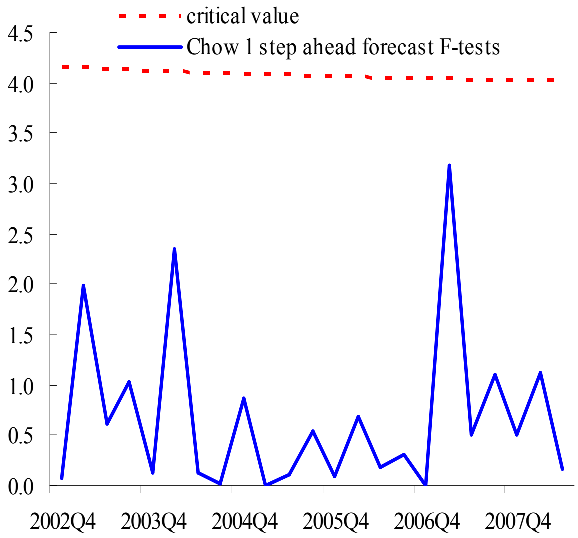
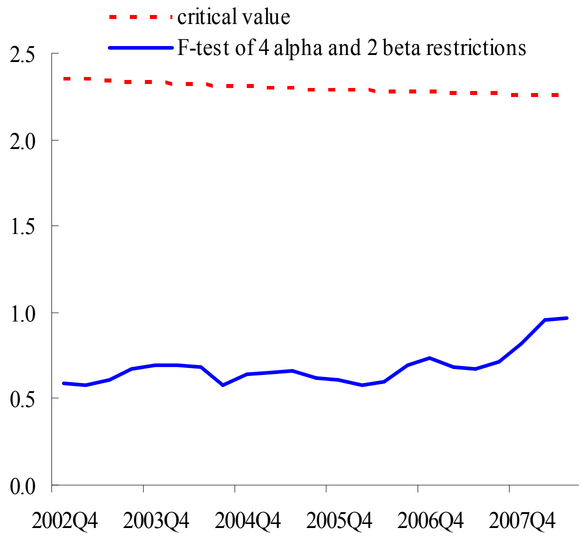


Figure 2b
Recursive F-test on alpha and beta restrictions



Turning to the cointegration relation obtained from model 3-r, it displays certain similarities with that obtained both for model 1-n, with regard to the large positive "peak" in 2000, as well as with that obtained for model 2-n with regard to the assessment of developments in the period 2004–2008. Contrary to both nominal models, model 3-r displays more frequent crossings of the average level. The cointegration relation derived from model 4-r contrasts significantly with that of the other three models. It displays several longer episodes of upward and downward movements. For the more recent period between 2004 and 2008, the assessment of money holdings relative to the long-run determinants would tend to confirm the results obtained from models 2-n and 3-r.

3.2.3. The adjustment to the long-run relationship – the αs

With regard to the variables involved in the adjustment to the long-run equilibrium, the tests for weak exogeneity of the variables (presented in Table 4) provide guidance for imposing the exclusion restriction on the a-vector in equation 3.

Table 6
The Loading Factors α

Model 1-n							Test
	Δm_t	Δpc_t	Δrc_t	$\Delta rthw_t$	$\Delta(blr_t - own_t)$	$\Delta GL1_t$	F(4,58)
	-0.0501 [0.009]	-	0.039 [0.009]	-	-	-	1.52 [0.21]
Model 2-n	Δm_t	Δpc_t	Δrc_t	Δrtw_t	$\Delta(blr_t - own_t)$	$\Delta(p\text{-}e)_t$	F(1,54)
	-0.111 [0.009]	-0.046 [0.014]	0.112 [0.022]	0.037 [0.016]	0.628 [0.203]	-	0.08 [0.79]
Model 3-r	$\Delta(m - pc)_t$	Δrdi_t	$\Delta\Delta pc_t$	$\Delta(IRL - IRS)_t$	$\Delta(p - e)_t$		F(3,62)
	-0.026 [0.006]	0.027 [0.006]	-	-	-		1.76 [0.16]
Model 4-r	$\Delta(m - pc)_t$	Δrc_t	$\Delta\Delta pc_t$	$\Delta(IRL - OWN)_t$	$\Delta(p - e)_t$	ΔC^e_t	F(4,61)
	-0.044 [0.009]	-	-	-	-	34.14 [7.85]	0.96 [0.44]

β restricted as in Table 5, standard errors in square brackets.

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The following restrictions on the load factors are compatible with the data:

1. In model 1-n, the test indicates that the load factors on the change in the price deflator, the change in wealth, the interest rate spread and the uncertainty measure can be restricted to zero. This leaves two variables to adjust to disequilibria, *m* and *rc*. The parameters for these two load factors are highly significant, with nominal money and real consumption helping to reduce the disequilibrium in the long-run relationship. A joint test for the restrictions placed on the α-vector cannot be rejected at conventional significance levels (see Table 6). This notwithstanding, the speed of adjustment observed for both variables is rather low, as commonly found in studies of household sector money demand.[18] This renders the short-run dynamics more important. Recursive estimation of the load factors indicates that the parameter estimate has remained unchanged between 2002 Q4 and 2008 Q3, while the same exercise for both *α*- and *ß*-restrictions shows a slight increase since mid-2007, while remaining well below the 5% significance threshold (see Figure 2b).
2. The price earnings ratio is the only weakly exogenous variable in the cointegrated VAR of model 2-n (see Table 4). The parameters for the four remaining load factors are highly significant (at the 5% significance level), with nominal money, real consumption and real wealth helping to reduce the disequilibrium, while again the opposite effect is exerted on the price level (see Table 6).[19] In real terms, however, money still equilibrium corrects.
3. In model 3-r, the tests on the weak exogeneity of the variables suggest that only the yield curve is weakly exogenous (see Table 4). Additional restrictions on the load factors for inflation and the price earnings ratio are not rejected. Therefore, only real money and real disposable income adjust to disequilibria (see Table 6). The speed of adjustment for both variables is highly significant, but very low.
4. The weak exogeneity tests in Table 4 indicate that only real household balances and expectations with regard to economic activity adjust to long-run disequilibria. The t-statistic on the load factor in the household M3 equation is 4.7 and well above the rule of thumb value provided by Juselius (2006). This supports the view that money holdings adjust to imbalances. The tests also suggest that the other variables (except the expectations) are pushing factors for monetary developments (see Table 6). This contrasts with the finding from model 1-n which indicated that real consumption adjusts to monetary disequilibria. However, in model 4-r consumer expectations adjust. Granger causality tests also provide weak evidence that money affects consumption expenditure in this model (p-value = 0.11), while an indirect effect is detected from money to consumer expectations, onto the price earnings ratio (p-value = 0.05) and finally on consumption (p-value = 0.06).

3.2.4. Models' explanatory power for household M3 and misspecification tests

The cointegrated VAR models 1-n and 2-n explain the quarterly changes in households' money balances well, with an adjusted R^2 of 0.69 and 0.73, respectively. The goodness of fit of the equation is also illustrated by Chart D in annex 1 which compares actual and fitted data. The residuals in both models for the household M3 equation show a large spike at the end of 2002 (see Chart E). The cointegrated VAR models incorporating real household balances, model 3-r and 4-r, explain the quarterly changes in households' M3 balances less well than the nominal models. The respective adjusted R^2 is 0.49 and 0.59. Model 2-n also fits the development in real consumption and total wealth surprisingly well with a respective adjusted R^2 of 0.51 and 0.92, while model 3-r is able to explain a noticeable share of the quarterly variation in real disposable income, as evidenced by an adjusted R^2 of 0.42. Model 4-r tracks the quarterly variation in consumer expectations quite well, with an adjusted R^2 of 0.41.

[18] See for instance von Landesberger (2007).

[19] A joint test for all restrictions imposed in model 2-n is not rejected at conventional significance levels [F(4,53) = 0.93, p-value = 0.46]

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Table 7
Residual properties for coinetgrated VAR

Specification test	Test statistic	p-value	Stability test	Test statistic	p-value
Model 1-n					
LM-AR(1)	F(36,51) = 0.85	0.69	LM-PC(3) vs. deterministic variables	1.56	0.67
LM-AR(4)	F(36,48) = 0.88	0.65	LM-PC(3) cointegration	0.90	0.83
Multivariate ARCH	F(441,60) = 1.13	0.28	sup Q(t|T)	1.42	0.70
Normality	F(12,56)= 1.17	0.32	mean Q(t|T)	0.73	0.52
Model 2-n					
LM-AR(1)	F(36,47) = 0.70	0.87	LM-PC(3) vs. deterministic variables	0.76	0.86
LM-AR(4)	F(36,44) = 0.63	0.92	LM-PC(3) vs. cointegration	0.94	0.82
Multivariate ARCH	F(441,53) = 1.05	0.44	sup Q(t|T)	4.54	0.15
Normality	F(12,52) = 2.19	0.03	mean Q(t|T)	0.98	0.29
Model 3-r					
LM-AR(1)	F(25,56) = 1.02	0.46	LM-PC(3) vs. deterministic variables	1.36	0.71
LM-AR(4)	F(25,53) = 1.09	0.39	LM-PC(3) vs. cointegration	1.39	0.71
Multivariate ARCH	F(225,62) = 1.20	0.20	sup Q(t|T)	1.75	0.65
Normality	F(10,60) = 0.38	0.95	mean Q(t|T)	0.85	0.48
Model 4-r					
LM-AR(1)	F(36,54) = 1.17	0.30	LM-PC(3) vs. deterministic variables	0.50	0.92
LM-AR(4)	F(36,51) = 0.69	0.88	LM-PC(3) vs. cointegration	0.52	0.91
Multivariate ARCH	F(441,61) = 1.08	0.37	Nyblom sup Q(t|T)	0.79	0.35
Normality	F(12,59) = 2.79	0.00	Nyblom mean Q(t|T)	0.33	0.30

Notes: LM-AR(1) and LM-AR(4) test statistic calculated as in Johansen (1996). ARCH test follows Warne (2009). Normality test as proposed by Doornik and Hansen (1994). LM-PC(3) tests are based on Teräsvirta (1998) calculated using a third order Taylor expansion. Nyblom sup Q(t|T) and mean Q(t|T) computed as in Hansen and Johansen (1999).

In order to assess the statistical properties of the models, Table 7 reports results from several standard misspecification tests on the residuals of the cointegrated VARs. The misspecification tests indicate absence of autocorrelation of residuals for all the models. Multivariate ARCH effects can also not be detected in the residuals. In the case of models 2-n and 4-r, however, the residuals are not normally distributed. The Nyblom tests conditional on the full sample estimates for the constant and the lagged endogenous parameters do not reveal any instability of the long-run parameters for any of the models. Finally, the LM-tests against the alternative of non-linearity in the deterministic variables or the cointegration parameters, which would capture gradual shifts, also do not suggest parameter non-constancy.[20]

[20] See Teräsvirta (1998). Teräsvirta and Eliasson (2001) investigate non-linearity in an error correction model of UK money demand.

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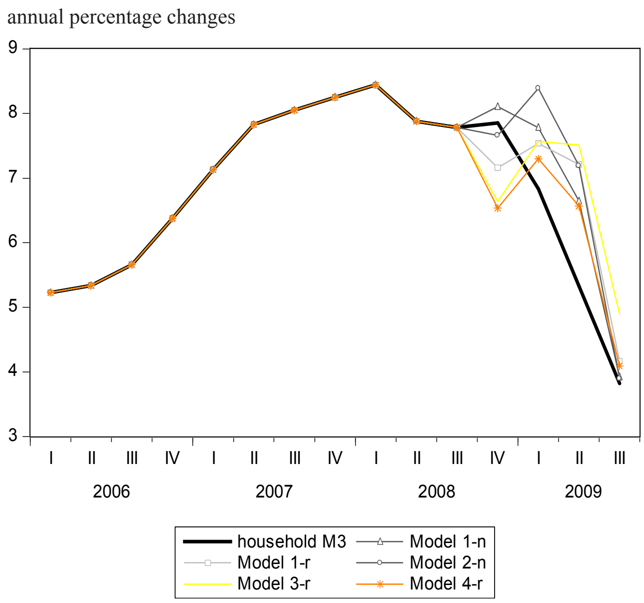
4. EVALUATING THE FORECASTING PERFORMANCE OF THE MONEY DEMAND SYSTEMS

The results presented in Section 3 suggest that the four models describe money demand by euro area households in a satisfactory manner, when judged, for instance, by the in-sample fit and standard misspecification tests. In addition, the estimates for the parameters allow for a theory-consistent interpretation and thereby support the view that money demand relationships have been identified. However, in order to gain additional insights on the models' ability to explain monetary developments, the last four available observations (2008 Q4 to 2009 Q3) are used to produce out-of-sample forecasts. The period covers the intensification of the financial turmoil following the default of Lehman Brothers, which has proven to be challenging for empirical models. In this context, it should be borne in mind that complicated models may have more explanatory power in sample, but also tend to include more variables that can lead to bad forecast results when changes to the economic environment occur. Thus, a more parsimonious specification may be advantageous.

Figure 3 illustrates the forecasts in terms of annual growth rates of household M3 using the actual observations for the other explanatory variables. It suggests that the nominal models and in particular model 1-n predicts monetary developments quite well, capturing the overall pattern of the slowdown, while the predictions from the real specifications suffer from the lack of capturing the rapid slowdown in price developments. More specifically, the strength of the slowdown in household M3 growth in 2009 Q2 was not captured by the models in a convincing manner, while most models do produce a prediction close to the actual outcome for 2009 Q3.

Figure 3
Forecast performance



Notes: for the period 2008 Q3–2009 Q3 actual observations for other explanatory variables.

In short forecasting samples characterized by structural breaks, cointegration models may not be able to exploit the advantage of having an identified long-run relationship. A different perspective on the ability of the cointegrated VAR systems to explain monetary developments is obtained when simulating out-of-sample the money growth for an extended period. Figure 4 shows the outcome of such an exercise conducted with model 2-n and 4-r from 1999 Q1 onward. The forecast for real household M3 growth settles at a stable annual steady state growth rate of 2.5% within eight quarters for model 4-r and takes twice as long for model 2-n. Investigating alternative

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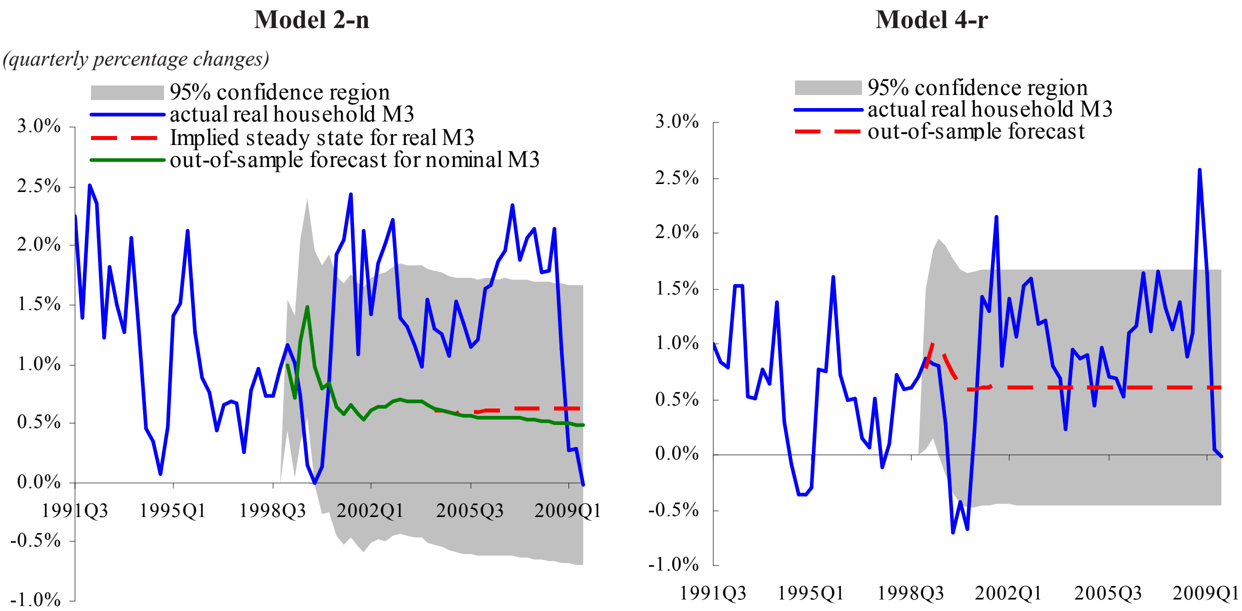
forecast horizons provides similar steady state growth rates. The simulations support the ECB's assessment that a number of exceptional shocks have hit euro area monetary developments over the past ten years, evidenced by the fact that actual M3 growth leaves the 95% confidence region: in 2000 in the context of the dot-com bubble with very low money growth, and later with very high money growth in 2001 Q3 owing to 11 September 2001 and in 2008 Q4 after the default of Lehman Brothers.[21] However, the models also clearly illustrate that household M3 growth has exhibited protracted periods of above steady-state growth, between 2002 and 2004 linked to exceptional portfolio shifts into money, and between 2006 and 2008, as money growth has been boosted by rapid growth of loans for house purchases in the euro area.

Figure 4
Out-of-sample forecast performance 1999:1–2009:3



5. CROSS-CHECKING WITH OTHER ESTIMATION METHODS

A repeatedly voiced observation with regard to the standard cointegrated VAR methodology is that the parameters of interest in the long-run relationship may be affected by the inclusion in the VAR set-up of less relevant variables. In order to cross-check the results obtained with the Johansen methodology, an alternative estimation is conducted using Fully Modified-OLS proposed by Phillips and Hansen (1990). This is a single equation regression with non-stationary variables. In the presence of several model variables affected by the long-run relationships, i.e. not all variables are weakly exogenous, the FM-OLS estimator will not be efficient as the move to the single equation neglects relevant information that could lead to a better point estimate. Nonetheless, if the residuals of the jointly error correcting variables are uncorrelated, this may be a restrictive yet informative exercise. For this purpose, the constrained specifications are re-estimated. The outcome of the exercise is provided in Table 8 showing the point estimates for the parameters of the cointegration relationship from the cointegrated VAR and the FM-OLS procedure.

21 European Central Bank (2005, 2007).

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Table 8
Estimation results with FM-OLS

	Model 1-n			Model 2-n			Model 3-r			Model 4-r		
Parameter estimate on	Coint VAR	FM-OLS	*Bootstrapped interval 10%*	Coint VAR	FM-OLS	*Bootstrapped interval 10%*	Coint VAR	FM-OLS	*Bootstrapped interval 10%*	Coint VAR	FM-OLS	*Bootstrapped interval 10%*
m	1 -	1	-	1	1	-	-	-	-	-	-	-
m-pc	-		-	-	-	-	1	1	-	1	1	-
pc	+1	+1	-	+1	+1	-	-	-	-	-	-	-
rc	+0.66 [0.04]	+0.50 [0.08]	*+0.95* *+0.48*	+0.5	+0.42 [0.10]	*0.58* *0.42*	-	-	-	+1	+1.14 [0.01]	-
rdi	-	-	-	-	-	-	+1.82 [0.16]	+1.12 [0.01]	*+3.47* *-0.67*	-	-	-
rtw	-	-	-	+0.50	+0.55 [0.08]	*0.58* *0.42*	-	-	-	-	-	-
rthw	+0.66 [0.04]	+0.50 [0.07]	*+0.95* *+0.48*	-	-	-	-	-	-	-	-	-
dpc	-	-	-	-	-	-	-0.23 [0.08]	-0.06 [0.04]	*+0.67* *-1.59*	0 -	-	-
blr-own	-0.67 [0.10]	-0.07 [0.04]	*-0.34* *-1.56*	-0.07 [0.03]	-0.06 [0.04]	*-0.01* *-0.16*	-	-	-	-	-	-
IRL-IRS	-	-	-	-	-	-	-0.07 [0.02]	-0.00 [0.01]	*+0.19* *-0.42*	-	-	-
IRL-OWN	-	-	-	-	-	-	-	-	-	-0.08 [0.02]	-0.04 [0.01]	*-0.05* *-0.16*
GL1	+1.20 [0.29]	-0.03 [0.17]	*+0.42* *+1.99*	-	-	-	-	-	-	-	-	-
p-e	-	-	-	-0.23 [0.02]	-0.07 [0.03]	*-0.18* *-0.29*	-0.51 [0.08]	-0.17 [0.04]	*+0.94* *-2.33*	-0.45 [0.08]	-0.18 [0.04]	*-0.27* *-0.92*
C_EXP	-	-	-	-	-	-	-	-	-	+0.007 [0.001]	0.00 [0.00]	*+0.00* *+0.02*
Equality of models F-Test			F(5,60) = 6.8580 [0.03]			F(5,54) = 8.1874 [0.00]			F(4,62) = 8.7712 [0.00]			F(5,61) = 6.7599 [0.02]

Note: Standard errors in square brackets below coefficients.

The results suggest that the individual point estimates obtained by both econometric techniques are fairly similar. In addition, the table also reports the interval obtained from bootstrapping the β-estimates and imposing a 10% confidence interval.[22] In this respect a number of points are worth mentioning:

[22] A higher confidence level such as 5% would have increased the width of the confidence bands significantly.

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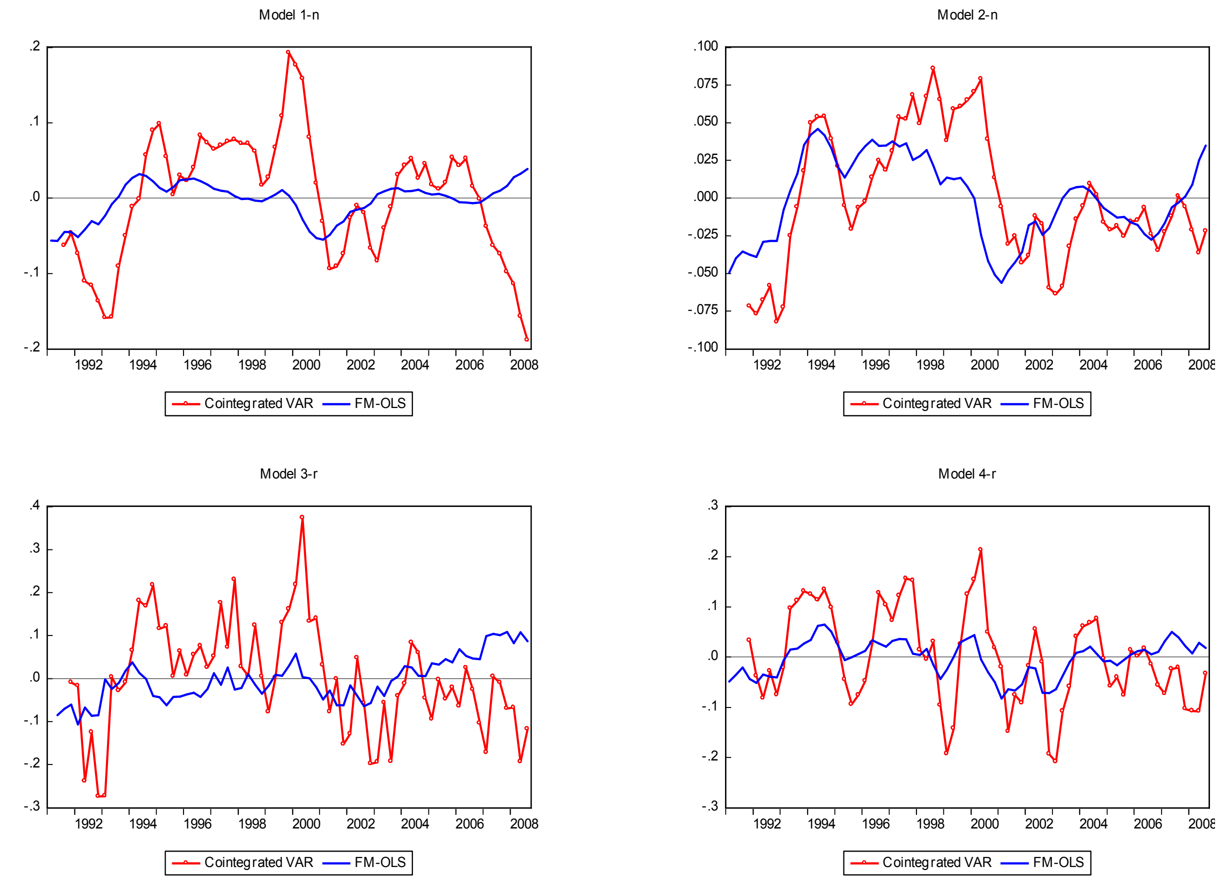
- The parameters of models 2 and 3 are included in the bootstrap interval, which are, however, excessively wide.
- The parameter estimates obtained with FM-OLS for the scale variables (consumption, disposable income, wealth) are generally included in the bootstrapped intervals of the cointegrated VAR.
- The outcome is more mixed for the other explanatory variables (opportunity costs, uncertainty related variables). The parameter estimates on these other variables, if significant in the FM-OLS equation, tend to be different from those obtained in the cointegrated VAR implying a distinctive assessment of the importance of these variables for money demand.
- Jointly imposing the parameter estimates obtained from the FM-OLS procedure in the original cointegrated VAR framework leads to a rejection of the equality of the estimates at the 5% level in all cases (see the last row with p-values in brackets).

However, gauging the similarity of the cointegration relationships by the cross correlation between their monetary overhang measures suggests that models 2-n and 4-r are quite similar in their assessment of actual money holdings relative to equilibrium household M3 holdings with contemporaneous correlation coefficients around 0.55 for the period 1991 Q1 to 2008 Q3. By contrast, for models 1-n and 3-r a similarity between the two measures cannot be found contemporaneously (0.19 and 0.00). Cross-correlation analysis suggests that the Johansen measures tend to lead their respective FMOLS measures with a slightly better match. Figure 5 shows the monetary overhangs from the cointegrated VAR (red line) and FM-OLS estimation (blue line) of all models.

Figure 5
Comparison of monetary overhang measures



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Monetary overhangs are often considered as an appropriate metric for measuring excess liquidity. When gauging the similarity of the assessment by whether both series are over or under the zero line in Figure 3, the superiority of models 2-n and 4-r is again clearly visible.

The chart also illustrates the different development in the estimated monetary imbalances during the financial turmoil from mid-2007 onwards. In this period a sharp decrease in the opportunity costs of holding money and an increase in uncertainty was observed. It is not surprising that the monetary overhang measures estimated using FM-OLS that downplay these factors differ especially in this period.

6. SUMMARY AND CONCLUSION

In the euro area, household holdings of M3 have been found to be informative with regard to developments in HICP inflation. An empirical framework permitting to analyse the driving factors for household money demand is therefore an important element for monetary analysis. The paper presented several different approaches to model the demand for nominal and real household M3 balances in the euro area. In investigating the long-run relationship between money, different scale variables and opportunity costs, only a few combinations may satisfy formal cointegration tests, even if an underlying cointegration relationship is present for a broader set of similar variables (see Ericsson, 1998). Several important outcomes have been found.

Neither nominal models rejects linear homogeneity between money balances and the price level. They therefore support the specification in real terms and suggest that in the long-run households are not subject to money illusion, in line with theoretical considerations.

1. Household money balances are never weakly exogenous with regard to the other variables of the cointegrated VARs and therefore always adjust to disequilibria between (real) money and its long-run determinants. That said, the models also provide evidence that the volume of transactions (proxied by real disposable income or real consumption) is affected by the monetary disequilibria and also adjusts. By contrast, measures of wealth, opportunity costs and financial market uncertainty are generally found to be the forces jointly determining the growth of money and real income/real consumption in the long-run.
2. In explaining households' broad money balances, wealth, and in particular housing wealth, is found to play an important role.[23] However, it seems to be wealth in conjunction with either real consumption expenditures or real disposable income that best captures households' notional level of money holdings. Omitting wealth from the specification leads to a sizeable increase in the income elasticity of money demand. At the same time, the inclusion of consumer expectations with regard to economic activity is able to offset this increase. This may reflect the fact that, in theory, wealth captures expectations on the future income path.
3. Interest rate developments seem to play a significant role for the development of household balances. Models specified with a double-log formulation for opportunity costs exhibit a markedly stronger impact than is the case for the semi-log functional forms. However, in all models, an increase in opportunity costs leads to a significant decline in (real) money holdings with the effect fully materialising after about 10 quarters.
4. The different models suggest that the impact of uncertainty on household balances is complex. Financial market uncertainty clearly stimulates M3. By contrast, economic uncertainty exhibits a more ambiguous impact on money holdings reflecting, on the one hand, the boosting impact stemming from the precautionary motive and, on the other, the dampening impact from a transactions motive.

[23] See de Bondt (2009) and Beyer (2009) for a similar finding with regard to aggregate M3.

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5. Correctly incorporating the persistent behaviour of interest rates and uncertainty into the money demand function is essential to adequately capture the driving forces impacting on money and expenditures as well as their mutual interaction. The different specifications presented suggest that there are several modelling approaches that can be undertaken.
6. All models pass a battery of misspecification and stability tests. Moreover, the parameter estimates are checked with a (probably more robust) single-equation approach. Especially at the end of the sample, differences in the models' estimates are obvious.

While the outcome of the exploration may not be seen as surprising as the estimates are consistent with results reported in the literature, the exercises presented help to better identify the determinants of euro area money demand and to interpret current monetary developments. As households' money demand captures the bulk of aggregate euro area M3, it should also be helpful in understanding the long-run money-price-nexus.

More generally, the exercise also provides insights that go beyond the portfolio allocation decision of households. According to our analysis, it is quite apparent that in equilibrium, households jointly determine consumption and broad money holdings, influenced by both wealth as well as interest rates. The importance of household money holdings for consumption expenditures may cast doubt on a purely passive role for money in this context. Moreover, as both bank lending rates and the own rate of households M3 are found significant, the determination of money holdings seems to interact with wealth and borrowing. In order to be able to analyse the interaction between money holdings, consumption and wealth more fully, the financing of households needs to be modelled as well, which goes beyond the scope of this paper.

References

Adalid, R. and C. Detken (2007) Liquidity Shocks and Asset Price Boom/Bust Cycles. ECB Working Paper 732.

Anderson, R. and S. Collins (1997) Modelling US Households' Demand for Liquid Wealth in an Era of Financial Change. *Journal of Money, Credit and Banking* 30, pp. 82–101.

Atta-Mensah, J. (2004a) The Demand for Money in a Stochastic Environment. Bank of Canada Working Paper 2004-7, March.

Atta-Mensah, J. (2004b) Money Demand and Economic Uncertainty. Bank of Canada Working Paper 2004-25, July.

Attanasio, O., L. Guiso and T. Jappelli (1998) The Demand for Money, Financial Innovation, and the Welfare Cost of Inflation: An Analysis with Household Data. National Bureau of Economic Research Working Paper 6593, June.

Beyer, A. (2009) A Stable Model for Euro Area Money Demand: Revisiting the Role of Wealth. ECB Working Paper No. 1111, November.

Boone, L. and P. van den Noord (2008) Wealth Effects on Money Demand in the Euro Area. *Empirical Economics* 34, pp. 525–536.

Brigden, A. and P. Mizen (1999) Money, Credit and Investment in the UK Corporate Sector. Bank of England Working Paper 100.

Butkiewicz, J.J. and M.M. McConnell (1995) The Stability of the Demand for Money and M1 Velocity: Evidence from the Sectoral Data. *Quarterly Review of Economics and Finance* 35/3, pp. 233–243.

Calza, A., D. Gerdesmeier and J. Levy (2001) Euro Area Money Demand: Measuring the Opportunity Costs Appropriately. IMF Working Paper 01/179, November.

Calza, A. And A. Zaghini (2010) Sectoral Money Demand and the Great Disinflation in the United States. *Journal of Money, Credit and Banking* 42, pp. 1663–1678.

Campbell, J.Y. and R.J. Shiller (1998) Valuation Ratios and the Long-run Stock Market Outlook. *Journal of Portfolio Management* 24, pp. 11–26.

Carstensen, K. (2006) Stock Market Downswing and the Stability of European Monetary Union Money Demand. *Journal of Business and Economic Statistics* 25, pp. 395–402.

Chadha, J., A. Haldane and N. Janssen (1998) Shoe Leather Costs Reconsidered, *Economic Journal* 108, pp. 363–382.

Chang, Y., J.Y. Park and K. Song (2002) Bootstrapping Cointegrating Regressions. Department of Economics Working Paper 2002-04, Rice University, February.

Chordia, T. and L. Shivakumar (2002) Momentum, Business Cycle, and Time-Varying Expected Returns, *Journal of Finance* 57, pp. 985–1019.

DOI: 10.7172/2353-6845.jbfe.2014.2.4

Coenen, G. and J.-L. Vega (2001) The Demand for M3 in the Euro Area. *Journal of Applied Econometrics* 16, pp. 727–748.
Cheung, Y.-W. and K.S. Lai (1993) Finite-Sample Sizes of Johansen's Likelihood Ratio Tests for Cointegration. *Oxford Bulletin of Economics and Statistics* 55, pp. 313–328.
Chrystal, A. and P. Mizen (2001) Consumption, Money and Lending: a joint model for the UK household sector. Bank of England Working Paper 134, May.
Davidson, R. and J.G. MacKinnon (1993) Estimation and Inference in Econometrics, New York, Oxford, Oxford University Press.
De Bondt, G. (2009) Euro Area Money Demand: Empirical evidence on the role of equity and labour markets. ECB Working Paper No. 1068, September.
Detken, C. and F. Smets (2004) Asset Price Booms and Monetary Policy. ECB Working Paper 364, May.
Doornik, J. and H. Hansen (1994) An Omnibus Test for Univariate and Multivariate Normality. Working Paper, Nuffield College, Oxford.
Ericsson, N.R. (1998) Empirical Modelling of Money Demand. *Empirical Economics* 23, pp. 295–315.
Ericsson, N.R., D.F. Hendry and H.-A. Tran (1994) Cointegration, Seasonality, Encompassing, and the Demand for Money in the United Kingdom, in: Hargreaves, C. (Ed.), Non-stationary Time-series Analysis and Cointegration, Oxford University Press, Oxford, pp. 179–224.
European Central Bank (2005) Money Demand and Uncertainty. Monthly Bulletin, October, pp. 57–73.
European Central Bank (2006a) Sectoral Money Holdings: Determinants and Recent Developments, Monthly Bulletin, August, pp. 59–72.
European Central Bank (2006b) Sectoral Money Holdings and the Information Content of Money with Respect to Inflation. Box 1, ECB Monthly Bulletin, September, pp. 18–20.
European Central Bank (2007) Interpreting Monetary Developments since mid-2004. Monthly Bulletin, July, pp. 51–74.
Fama, E. and K. French (2002) The Equity Premium. *The Journal of Finance* 56, pp. 637–659.
Feiss, N. and R. MacDonald (2001) The Instability of the Money Demand Function: an I(2) interpretation. *Oxford Bulletin of Economics and Statistics* 63, pp. 475–495.
Friedman, M. (1956) Studies in the Quantity Theory of Money. Chicago, University of Chicago Press.
Friedman, M. (1988) Money and the Stock Market. *Journal of Political Economy* 96, pp. 221–245.
Fujiki, H. and C. Hsiao (2008) Aggregate and Household Demand for Money: Evidence from Public Opinion Survey on Household Financial Assets and Liabilities. Bank of Japan Institute of Monetary and Economic Studies Discussion Working Paper No. 2008-E-17.
Garver, C. and L. Radecki (1987) The Household Demand for Money: Estimates from Cross-sectional Data. *Federal Reserve Bank of New York Quarterly Review*, Spring, pp. 29–34.
Gerdesmeier, D. (1996) Die Rolle des Vermögens in der Geldnachfrage. Discussion paper 96/5, Volkswirtschaftliche Forschungsgruppe der Deutschen Bundesbank, May.
Goldfeld, S. (1973) The Demand for Money Revisited. *Brookings Papers on Economic Activity* 1973/3, pp. 577–646.
Greiber, C. and W. Lemke (2005) Money Demand and Macroeconomic Uncertainty. Deutsche Bundesbank Discussion Paper 26-2005.
Greiber, C. and R. Setzer (2007) Money and Housing: Evidence for the Euro Area and the US. Deutsche Bundesbank Discussion Paper 12-2007.
Hansen, H. and S. Johansen (1999) Some Tests for Parameter Constancy in Cointegrated VAR-models. *Econometrics Journal* 2, pp. 306–333.
Jain, P. and C.-G. Moon (1994) Sectoral Money Demand: A cointegration approach. *Review of Economics and Statistics* 76, pp. 196–202.
Johansen, S. (1996) Likelihood-Based Inference in Cointegrated Vector Autoregressive Models. 2nd ed., Oxford, Oxford University Press.
Juselius, K. (2006) The Cointegrated VAR Model: Methodology and Applications. Oxford, Oxford University Press.
Kontolemis, Z.G. (2002) Money Demand in the Euro Area: Where Do We Stand (Today)? IMF Working Paper 02/185, November.
Kilian, L. (2001) Impulse Response Analysis in Vector Autoregressions with Unkown Lag Order. *Journal of Forecasting* 20, pp. 161–179.
Kwiatkowski, D., P.C.B. Phillips, P. Schmidt and Y. Shin (1992) Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root. *Journal of Econometrics* 54, pp. 159–178.
Krupkina, A., Ponomarenko, A. (2013): Money demand models for Russia: A sectoral approach, Bank of Finland Institute for Economies in Transition Discussion Papers 31/2013.
Laumas, G.S. (1979) The Stability of the Demand for Money by the Household Sector: A Note. *Southern Economic Journal* 46, pp. 603–608.
Lucas, R. (2000) Inflation and Welfare. *Econometrica* 68, pp. 247–274.

DOI: 10.7172/2353-6845.jbfe.2014.2.4

Lütkepohl, H. and P. Saikkonen (1997) Order Selection in Testing for the Cointegrating Rank of a VAR Process. Sonderforschungsbereich 373: Quantification and Simulation of Economic Processes 93 (SFB 373 Papers).

MacKinnon, J.G. (1996) Numerical Distribution Functions for Unit Root and Cointegration Tests. *Journal of Applied Econometrics* 11, pp. 601–618.

MacKinnon, J.G., A.A. Haug and L. Michelis (1999) Numerical Distribution Functions of Likelihood Ratio Tests for Cointegration. *Journal of Applied Econometrics* 14, pp. 563–577.

Martinez-Carrascal, C. and J. von Landesberger (2010) Explaining the Money Demand of Non-Financial Corporations in the Euro Area: A Macro and a Micro View. ECB Working Paper No. 1257.

Muraru, A., Ruxandaa, G. (2011): Household money demand in Romania. Evidence from cointegrated VAR. *Technological and Economic Development of Economy* 17, pp. 382–396.

Park, J.Y. (2003) Bootstrap Unit Root Tests. *Econometrica* 71, pp. 1845–1895.

Petursson, T. (2000) The Representative Household's Demand for Money in a Cointegrated VAR Model. *Econometrics Journal* 3, pp. 162–176.

Phillips, P.C.B. and B.E. Hansen (1990) Statistical Inference in Instrumental Variables Regression with I(1) Processes. *Review of Economic Studies* 57, pp. 99–125.

Read, V. (1996) Sectoral Disaggregation of German M3. Deutsche Bundesbank Discussion Paper 1-1996.

Reimers, H.-E. (1991) Comparisons of Tests for Multivariate Cointegration. *Statistical Papers* 33, pp. 335–359.

Stracca, L. (2001) The Functional Form of the Demand for Euro Area M1. *The Manchester School* 71, pp. 172–204.

Stern, L. and M.L. Stern (2008) Expected Equity Returns and the Demand for Money. *The B.E. Journal of Macroeconomics* 8, Iss. 1, Art 18.

Swensen, A.R. (2006) Bootstrap Algorithms for Testing and Determining the Cointegration Rank in VAR Models. *Econometrica* 74, pp. 1699–1714.

Teräsvirta, T. (1998) Modelling Economic Relationships with Smooth Transition Regressions, in: Ullah A, Giles, D. E. A. (Eds.), Handbook of Applied Economic Statistics, New York et al., Dekker, pp. 507–552.

Teräsvirta, T. and A.C. Eliasson (2001) Non-linear Error Correction and the UK Demand for Broad Money, 1878–1993. *Journal of Applied Econometrics* 16, pp. 277–288.

Thomas, R. (1997a) The Demand for M4: A Sectoral Analysis Part 1 – The Personal Sector. Bank of England Working Paper 61.

Tin, J. (2008) An Empirical Examination of the Inventory – Theoretic Model of Precautionary Money Demand. *Economics Letters* 99, pp. 204–205.

Tobin, J. (1969) A General Equilibrium Approach to Monetary Theory. *Journal of Money, Credit and Banking* 1, pp. 15–29.

Von Landesberger, J. (2007) Sectoral Money Demand Models for the Euro Area Based on a Common Set of Determinants. ECB Working Paper No. 741, March.

Warne, A. (2006) Bayesian Inference in Cointegrated VAR Models with Applications to the Demand for Euro Area M3. ECB Working Paper No. 692, November.

Warne, A. (2008) Structural VAR, version 0.50, release 1.

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ANNEX 1

Table 9
Results of unit root tests

Variables	ADF (D,X)	ADF t-Statistic	ADF p-value*	KPSS	KPSS LM-Statistic	critical value 10%	critical value 5%
level						**10%**	**5%**
m	(CT,6)	-0.98	0.94	(CT)	0.6938	0.119	0.146
pc	(CT,9)	-1.14	0.92	(CT)	0.1049	0.119	0.146
rtw	(CT,12)	-1.75	0.72	(CT)	0.2358	0.119	0.146
rthw	(CT,12)	-1.88	0.65	(CT)	0.1908	0.119	0.146
rc	(CT,10)	-0.84	0.96	(CT)	0.1349	0.119	0.146
rdi	(CT,11)	-1.93	0.63	(CT)	0.1006	0.119	0.146
C^e	(C,11)	-2.19	0.21	(C)	0.1999	0.347	0.463
UN^e	(C,4)	-3.03	0.04	(C)	0.3467	0.347	0.463
GL1	(C,12)	-0.88	0.79	(C)	0.3903	0.347	0.463
p-e	(C,3)	-1.33	0.61	(C)	0.1497	0.347	0.463
IRL-IRS	(C,11)	-3.95	0.00	(C)	0.2156	0.347	0.463
blr-own	(C,1)	-1.46	0.55	(C)	0.3465	0.347	0.463
IRL-OWN	(C,10)	-2.61	0.10	(C)	0.4503	0.347	0.463
1st difference							
m	(C,0)	-3.45	0.01	(C)	0.2072	0.347	0.463
pc	(C,2)	-3.55	0.01	(C)	0.3393	0.347	0.463
rtw	(C,3)	-2.09	0.25	(C)	0.1928	0.347	0.463
rthw	(C,11)	-2.79	0.07	(C)	0.3706	0.347	0.463
rc	(C,9)	-2.30	0.17	(C)	0.1827	0.347	0.463
rdi	(C,10)	-1.94	0.31	(C)	0.1834	0.347	0.463
C^e	(C,10)	-3.40	0.00	(C)	0.1232	0.347	0.463
UN^e	(C,4)	-3.64	0.00	(C)	0.0510	0.347	0.463
GL1	(N,11)	-3.88	0.00	(C)	0.0882	0.119	0.146
p-e	(N,0)	-6.87	0.00	(C)	0.3404	0.347	0.463
IRL-IRS	(N,0)	-5.54	0.00	(C)	0.1625	0.347	0.463
blr-own	(N,2)	-4.70	0.00	(C)	0.2546	0.347	0.463
IRL-OWN	(N,9)	-4.18	0.00	(C)	0.0465	0.347	0.463

Note: ADF-test: with MacKinnon (1996) one-sided p-values, KPSS: Kwiatkowski-Phillips-Schmidt-Shin (1992, Table 1). (D,X) with D indicating that the estimated regression includes the following deterministic terms: C – constant, CT – constant and trend, N – no deterministic terms. X indicates the number of lagged endogenous terms retained in the estimated test regression (with at least 5% significance) starting from a maximum of 12 lags. Cut-off is determined by sequential testing on the t-statistic of the lagged endogenous variables with at least 5% significance level. KPSS test using Bartlett kernel with cut-off determined by automatic Andrews (1991) procedure.

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Table 10
Lag length determination for VAR

Model	Criterion	Lag length				
		0	1	2	3	4
1-n	Likelihood Ratio Test	NA	2186.05	119.68*	43.86	43.16
	Akaike Information Criterion	-24.03	-44.46	-45.74*	-45.58	45.39
2-n	Likelihood Ratio Test	NA	1330.82	217.75	109.98*	31.22
	Akaike Information Criterion	-19.00	-40.11	-43.07	-44.29*	-43.96
3-r	Likelihood Ratio Test	NA	1484.53	55.73*	28.34	22.46
	Akaike Information Criterion	-3.21	-27.20	-27.46*	-27.27	-27.01
4-r	Likelihood Ratio Test	NA	937.35	74.14*	36.97	29.26
	Akaike Information Criterion	-7.34	-22.13	-22.45*	-22.14	-21.76

Table 11
Tests for stationarity of variables

	Null Hypothesis: variable k is stationary Alternative hypothesis: variable k is not stationary							
	Model 1-n		**Model 2-n**		**Model 3-r**		**Model 4-r**	
Equation for	F(5,59)	p-value	F(5,53)	p-value	F(1,60)	p-value	F(5,60)	p-value
m	7.49	0.03	9.68	0.00	-	-	-	-
m-pc	-	-	-	-	8.86	0.01	7.98	0.01
pc	7.44	0.03	9.40	0.00	-	-	-	-
rc	7.64	0.03	8.99	0.00	-	-	7.55	0.02
rdi	-	-	-	-	8.34	0.01	-	-
rtw	-	-	9.38	0.00	-	-	-	-
rthw	7.43	0.05	-	-	-	-	-	-
dpc	-	-	-	-	4.78	0.03	5.08	0.03
blr-own	6.51	0.04	8.30	0.00	-	-	-	-
IRL-IRS	-	-	-	-	6.97	0.01	-	-
IRL-OWN	-	-	-	-	-	-	7.30	0.01
GL1	7.42	0.02	-	-	-	-	-	-
p-e	-	-	7.55	0.00	5.47	0.02	5.13	0.03
C^e	-	-	-	-	-	-	3.18	0.11

Note: p-values derived from comparison with respective bootstrapped distributions.

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Chart A
Main variables used in the cointegrated VAR systems

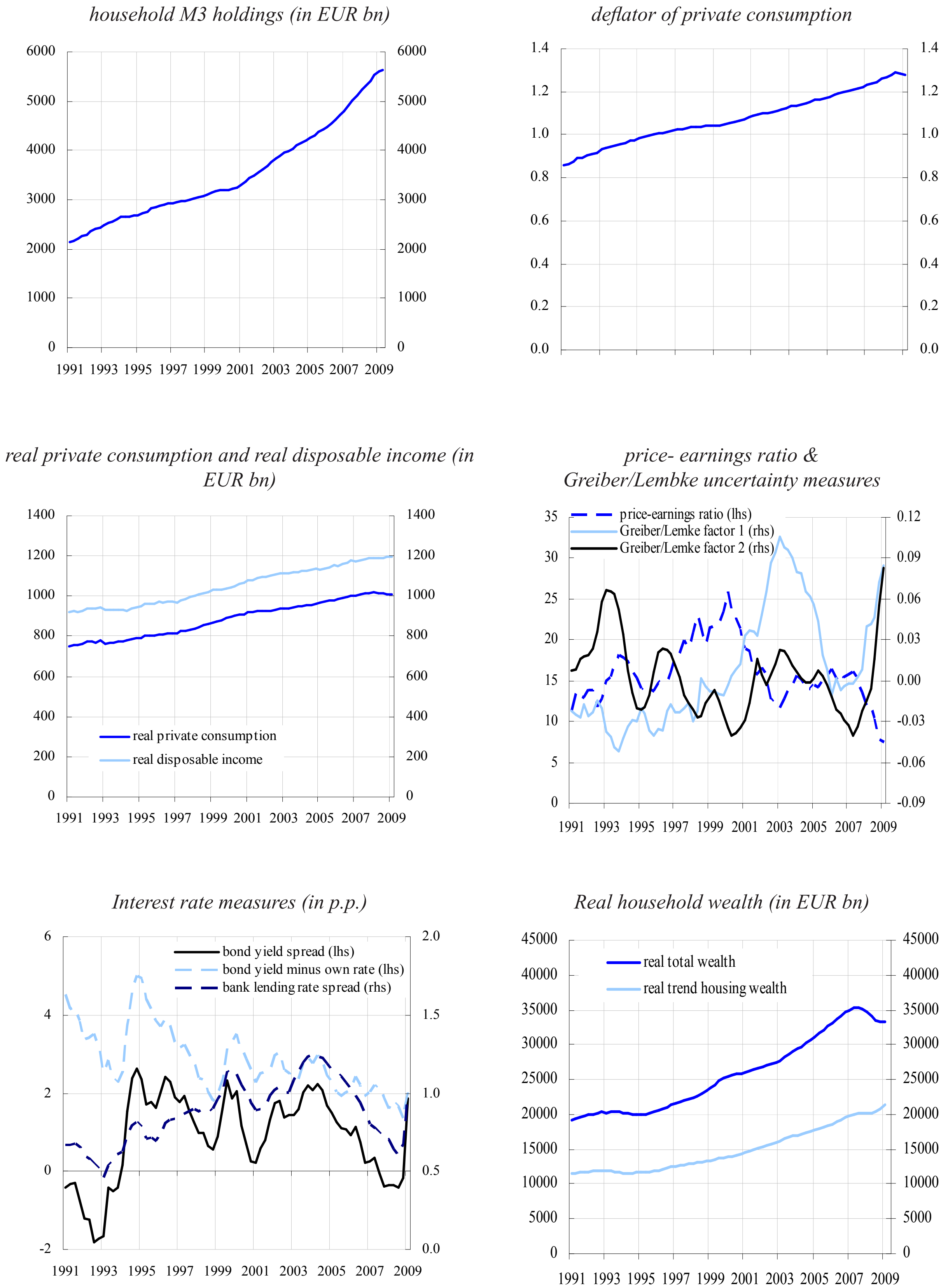

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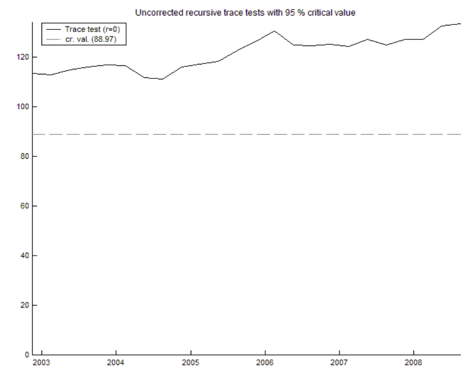
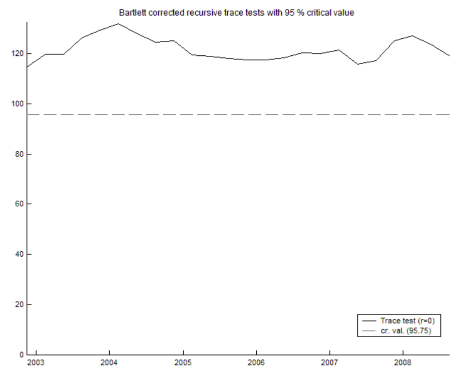
Chart B
Recursive trace test results

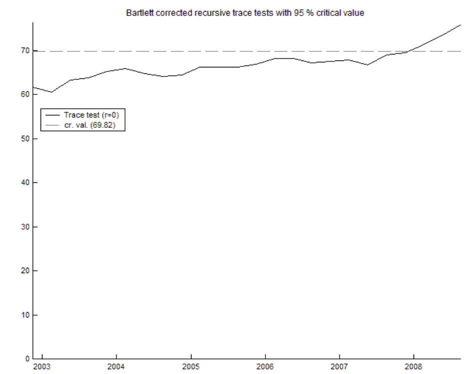
Model 1-n



Model 2-n



Model 3-r



Model 4-r

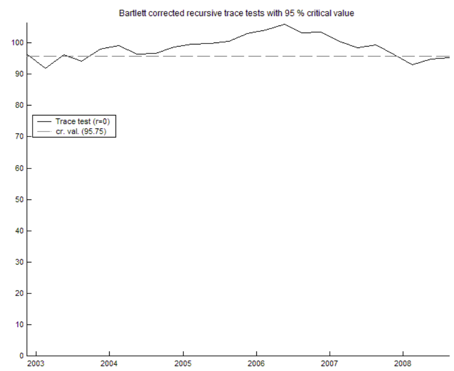
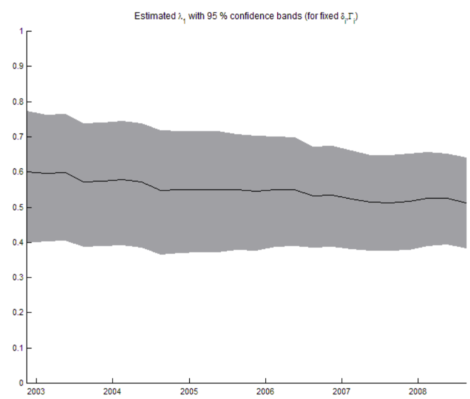
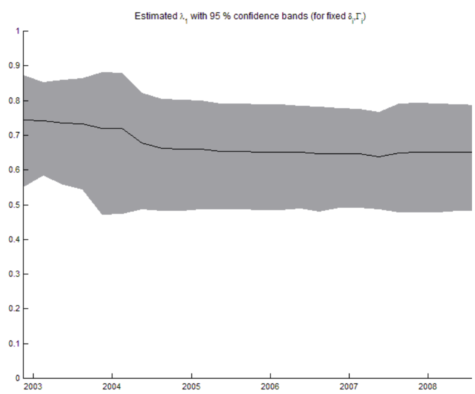


Chart C
Estimated λ_1 with 95% confidence bands (for fixed short run dynamics)

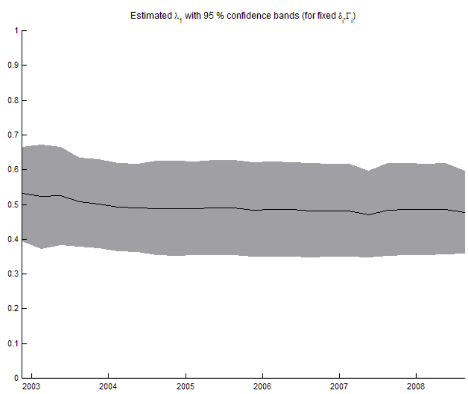
Model 1-n



Model 2-n



Model 3-r



Model 4-r

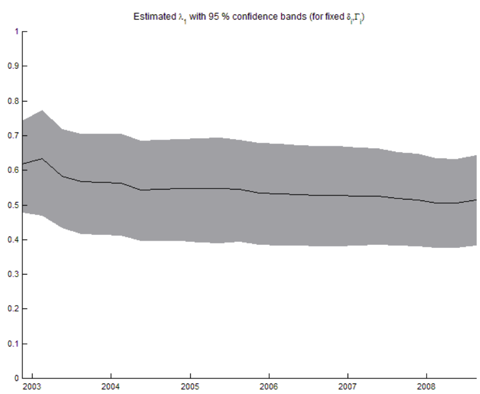


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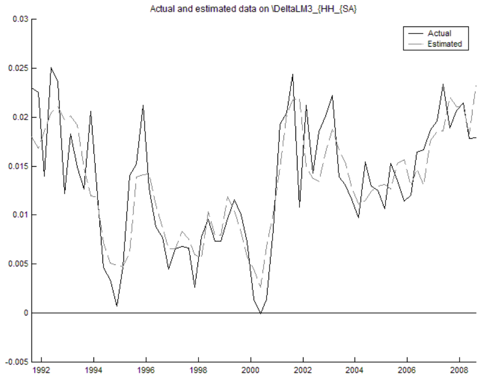
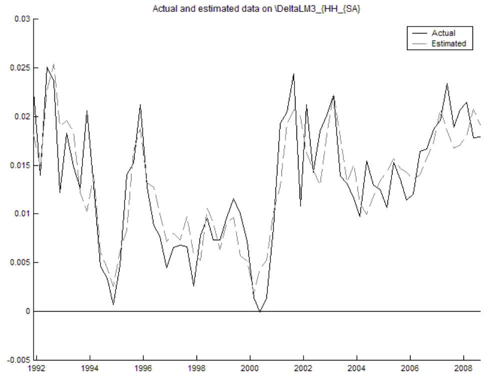
Chart D
Actual & fitted

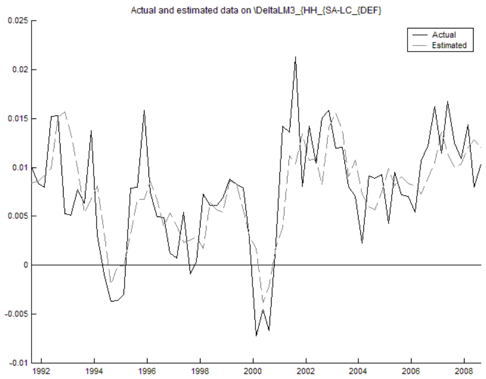
1-n



2-n



3-r



4-r

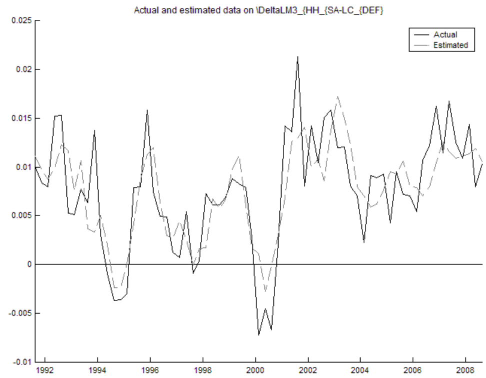
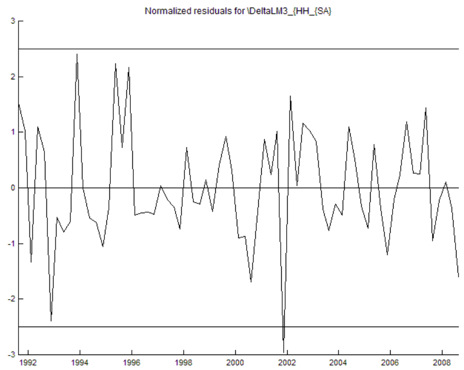
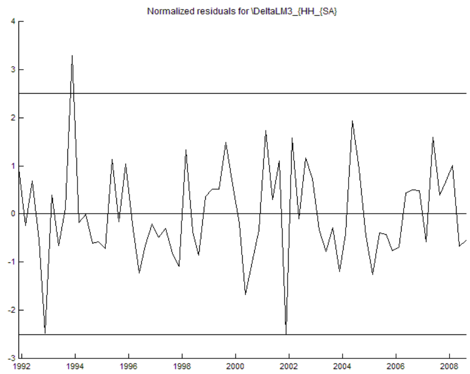


Chart E
Normalised residuals

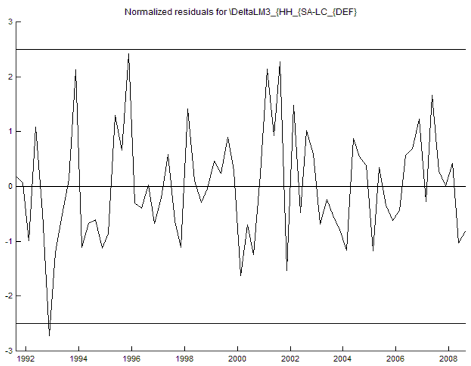
1-n



2-n



3-r



4-r

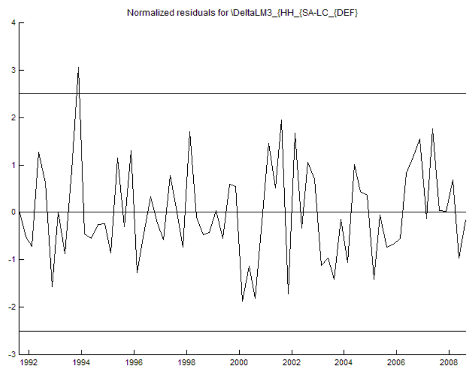


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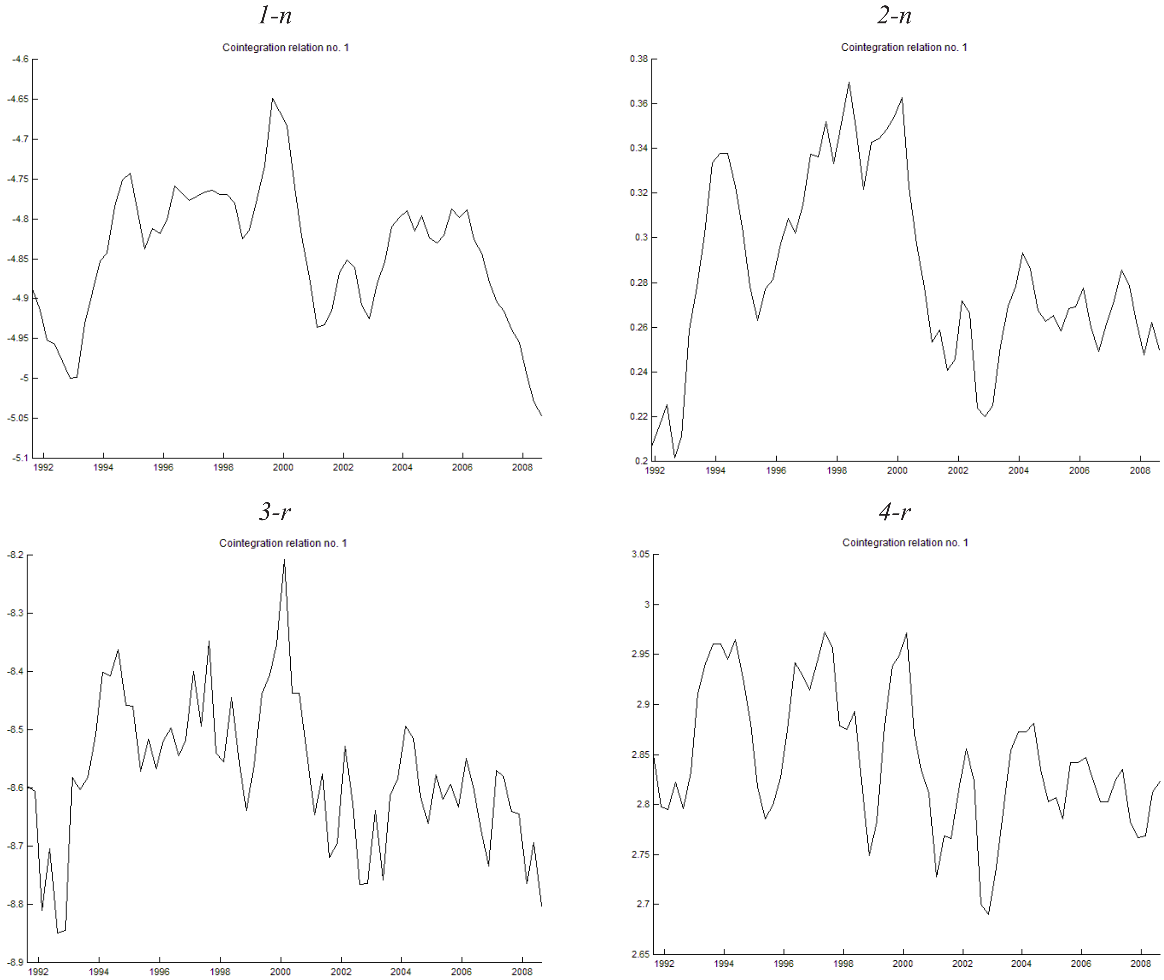
Chart F
Cointegration relationships



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ANNEX 2
Model 1-r

Table 12
Lag length selection for model 1-r

Lag length	**0**	**1**	**2**	**3**	**4**
Likelihood Ratio Test	NA	1089.94	125.62*	30.85	48.57
Akaike Information Criterion	-26.48	-44.18	-45.51*	-45.08	45.21

Table 13
Residual properties for model 1-r

Model 1-r					
LM-AR(1)	F(36,48) = 1.09	0.38	Multivariate ARCH	F(441,61) = 1.19	0.20
LM-AR(4)	F(36,51) = 1.14	0.33	Normality	F(12,59) = 0.85	0.60
For rank 1					
Nyblom Sup F		18.97			0.46
Nyblom Mean		5.49			0.50
PKK S(9)		1.32			0.39

Table 14
Trace test results for model 1-r

1-r	test	124.59	67.98	39.25	13.97	6.20	1.44	-
	p-value	0.000	0.000	0.259	/	/	0.230	-
	bootstrapped p-value	0.048	0.619	0.717	0.950	0.657	0.449	-

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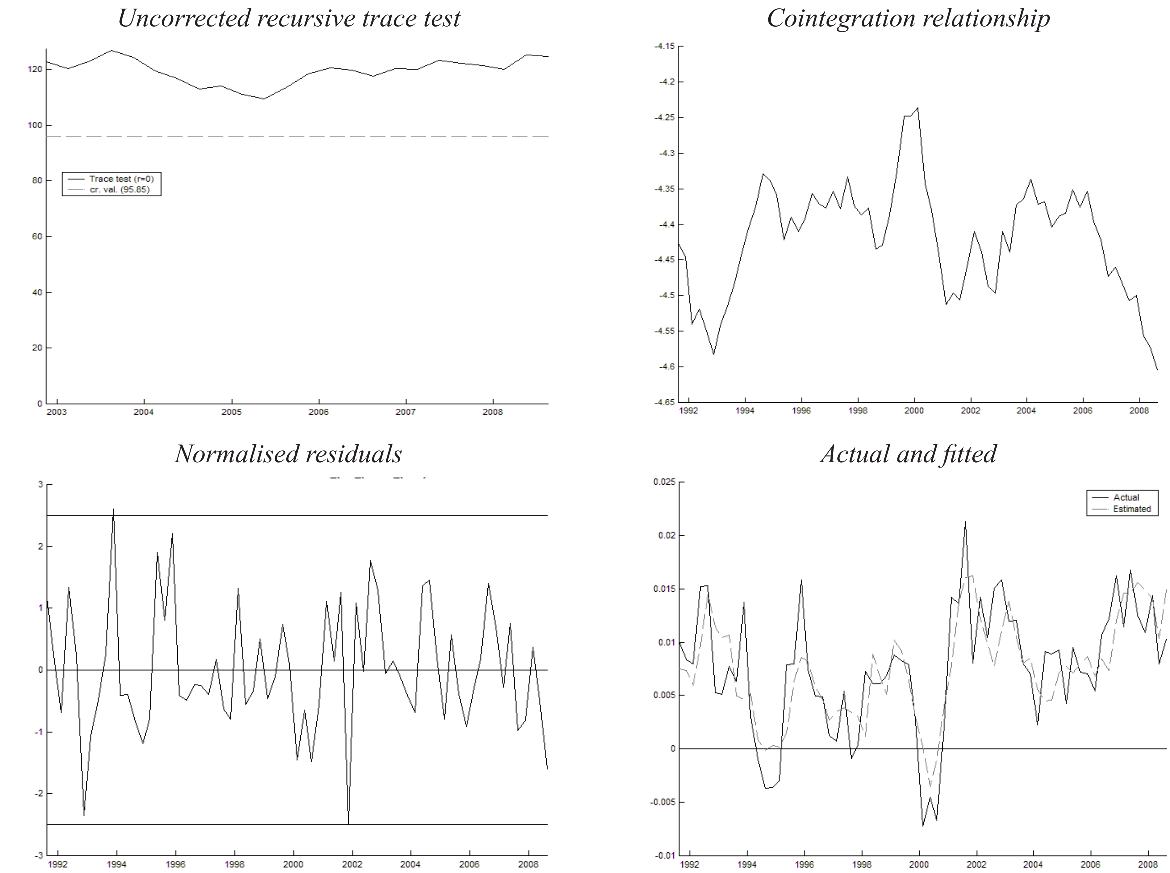
Chart G
Charts for model 1-r



Long-run relationship for model 1-r

$$\begin{bmatrix} \Delta(m-pc)_t \\ \Delta rc_t \\ \Delta rthw_t \\ dpc_t \\ \Delta(blr_t - own_t) \\ \Delta GL1_t \end{bmatrix} = \begin{bmatrix} \underset{0.008}{-0.044} \\ - \\ \underset{0.009}{0.038} \\ \underset{0.005}{0.011} \\ - \\ - \end{bmatrix} \left[(m-pc)_{t-1} - 0.50 rc_{t-1} - \underset{0.04}{0.74} rthw_{t-1} + \underset{0.09}{0.77} (blr - own)_{t-1} - \underset{0.22}{1.29} GL1_{t-1} + \underset{3.01}{7.00} dpc_{t-1} \right] + \ldots$$

Table 15
Residual properties for Model 1-r

Specification test	Test statistic	p-value	Stability tests	Test statistic	p-value
LM-AR(1)	F(36,53) = 0.76	0.81	LM-PC(18) lagged endogenous	10.51	0.91
LM-AR(4)	F(36,50) = 0.68	0.89	LM-PC(3) cointegration	1.22	0.75
Multivariate ARCH	F(441,60) = 1.09	0.35	Nyblom sup Q(t|T)	1.9315	0.91
Normality	F(12,58) = 0.58	0.85	Nyblom mean Q(t|T)	0.9648	0.82

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