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Still Valid at the Zero Lower Bound

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Re-vitalizing Money Demand in the Euro Area: Still Valid at the Zero Lower Bound

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Abstract

The analysis of monetary developments have always been a cornerstone of the ECB's monetary analysis and, thus, of its overall monetary policy strategy. In this respect, money demand models provide a framework for explaining monetary developments and assessing price stability over the medium term. It is a well-documented fact in the literature that, when interest rates are at the zero lower bound, the analysis of money stocks become even more important for monetary policy. Therefore, this paper re-investigates the stability properties of M3 demand in the euro area in the light of the recent economic crisis. A cointegration analysis is performed over the sample period 1983 Q1 and 2015 Q1 and leads to a well-identified model comprising real money balances, income, the long term interest rate and the own rate of M3 holdings. The specification appears to be robust against the Lucas critique of a policy dependent parameter regime, in the sense that no signs of breaks can be found when interest rates reach the zero lower bound. Furthermore, deviations of M3 from its equilibrium level do not point to substantial inflation pressure at the end of the sample. Excess liquidity models turn out to outperform the autoregressive benchmark, as they deliver more accurate CPI inflation forecasts, especially at the longer horizons. The inclusion of unconventional monetary policy measures does not contradict these findings.

Keywords: Euro area money demand, inflation forecasts, unconventional monetary policy

JEL-classification: E41, E44, E52, G11, G15

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Non-technical summary

The mandate of most central banks all over the world is to ensure low and stable inflation. At the same time, it is a widely accepted fact in the economic profession that, over medium and longer term horizons, inflation is essentially a monetary phenomenon. For this reason, the ECB's monetary policy strategy has been based on two pillars, commonly labelled as "economic analysis" and "monetary analysis". In the context of monetary analysis, money demand models play an important role for (at least) three important reasons. First, from a positive perspective, they provide a framework to explain monetary developments on the basis of the evolution of other macroeconomic variables, notably the price level, economic activity and interest rates. Second, from a normative or policy point of view, they can be used for determining the stock of money holdings or rate of monetary growth that is consistent with price stability over the medium term. Third, such benchmarks can be a useful starting point for constructing indicators relevant for taking monetary policy decisions.

Since its start, the ECB acknowledged the central character of the stability of the money/price relationship in the analytical framework for monetary analysis and this stability is typically assessed in the context of a money demand function.

This paper adds to the literature by addressing three key questions, namely: (i) whether, at the current juncture, there is still evidence of a stable long-run money demand relationship for the euro area even in a regime where interest rates seem to have reached a lower bound and in the context of the several repercussions of the financial crisis-- in regulatory and behavioural terms - for money-holders; (ii) whether broad monetary aggregates can help in predicting consumer price inflation and if so, over which time horizons; and (iii) to which extent the recent non-standard monetary policy measures have contributed to an outlook of higher inflation rates.

We find strong evidence of the presence of a stable, well-identified long-run money demand even in the most recent period. In contrast to the common view, however, the relationship does not show any signs of instability in periods where interest rates have reached the zero lower bound. Deviations of the money demand relationship from its equilibrium do not indicate substantial inflation pressure at the end of the sample. On a more general basis, excess liquidity models can outperform the autoregressive benchmark, as they deliver more accurate consumer inflation forecasts, especially at the longer horizons. The inclusion of unconventional monetary policy measures does not change these findings.

1 Introduction

The mandate of most central banks all over the world is to ensure low and stable inflation. In particular, at medium and longer term horizons, inflation is inherently a monetary phenomenon. Since money represents the unit of account, monetary developments are integral to the determination of prices and inflation.

As a consequence, it is a well-documented fact in the literature that the ECB's monetary policy strategy is based on two pillars, commonly labelled as "economic analysis" and "monetary analysis".[1] In the context of monetary analysis, money demand models play an important role for (at least) three important reasons. First, from a positive perspective, they provide a framework to explain monetary developments on the basis of the evolution of other macroeconomic variables, notably the price level, economic activity and interest rates and can, therefore, be seen as complementing the information coming from the economic analysis.

Second, from a normative or policy point of view, they can be used for determining the stock of money holdings or rate of monetary growth that is consistent with price stability over the medium term. In this context, the broad aggregate M3 contains information relevant for the assessment of risks to price stability in the medium to longer-term.

Third, such benchmarks can be a useful starting point for constructing indicators relevant for taking monetary policy decision (see Chapter 3 in Papademos and Stark (2010)). More precisely, in order to make an assessment of the monetary situation, some measure of the money stock needs to be examined against the background of the macroeconomic environment. The latter thus constitutes a natural benchmark against which actual monetary developments can be assessed.

Changes in the money stock could be driven by two distinct forces. Money might react to macroeconomic fundamentals or to conditions specific to the situation at hand. Further insights can be gained from decomposing money growth into the explanatory variables, such as output,

1 See, for instance, Chapter 2 in Papademos and Stark (2010).

prices and interest rates. Such an exercise is model specific, as the contributions will depend on which explanatory variables are considered in the model and the estimated parameters. The unexplained component of monetary growth should be rather minor, if the underlying specification is correct.

Money demand models have been a key tenet to the ECB's conduct of monetary analysis from the very beginning and, in outlining its monetary policy strategy in January 1999, the ECB acknowledged the central character of the stability of the money/price relationship in the analytical framework for monetary analysis. Indeed, the stability of this relationship is typically assessed in the context of a money demand function. The latter can be assessed using a variety of institutional analyses and statistical techniques. As regards the M3 demand models developed for the euro area, there has been evidence that some of them have become unstable over the years. Indeed, experience suggests that instabilities may be inevitable in an environment of perpetual structural change in an innovative financial system as well as at times financial crisis following the repercussions -- in regulatory and behavioural terms - for financial institutions and money-holders. As an example, more recently the relationship between money and its determinants may have been further distorted by the response of central banks to the recent financial crisis, as to provide stimulus for the real economy, interest rates were rapidly set to the zero lower bound, thus affecting the portfolio allocation decision of economic agents (see Reis, 2013). In addition, the introduction of unconventional monetary policy measures have most probably constituted a major shift in monetary policy.[2] As a consequence, according to the Lucas critique, money demand parameters can be expected to change..

Notwithstanding these arguments, the existence of a stable money demand function cannot be ruled out in advance from *a priori* point of view. Indeed, interest rates that are fixed at the zero lower bound limit the ability of a central bank to conduct monetary policy. However, in this situation, the use of other monetary policy instruments renders the analysis of the money

[2] See Williams (2014) for a more extensive overview on monetary policy at the zero lower bound (ZLB).

stock even more important. For this reason, this paper re-investigates the stability of M3 demand in the euro area, also covering the most recent period. The basic specification of the money demand function includes, as explanatory variables, real income, nominal long-term interest rates and the own rate of money holdings, with the opportunity cost variable being proxied by the term spread.

A number of striking conclusions emerge from the analysis. First, a long run money demand relationship can be identified for the euro area, which is characterized by stable parameters. While the income elasticity is not different from 1.0, the semi-elasticity of money with respect to the spread is around -0.1. These elasticities are robust to alternative estimation methods, and the corresponding error correction model survives a battery of specification tests. While still uncontested from a theoretical perspective, the Lucas critique seems to be not of relevance from an empirical point of view, as the regressors are superexogenous for real money balances. Second, the long-run relationship is employed to derive excess liquidity measures. Despite the accommodative monetary policy stance, especially in the most recent period, no evidence of excess liquidity can be found at the end of the sample. This result shows that the monetary transmission process is still impaired, as the impact on the monetary base (M0) is not yet translated further into the broader aggregate (M3). Finally, the model is used to predict HICP inflation. In particular, models based on excess liquidity measures outperform the autoregressive benchmark especially at longer forecasting horizons. The inclusion of variables related to unconventional monetary policies does not add much to the results.

The paper is structured as follows. Section 2 reviews previous studies on money demand in the euro area. Section 3 discusses the specification of money demand as well as the data. Section 4 holds the cointegration evidence of a long run relationship between money and its fundamentals, as well as tests on exogeneity. Section 5 presents the forecasting exercise. Finally, Section 6 concludes.

2 Previous evidence on money demand

Since the establishment of the euro area, a number of money demand models have been developed. To begin with, Fagan and Henry (1998) reported supportive evidence for a stable M3 demand, using income and interest rates as explanatory variables. Subsequent work by Coenen and Vega (2001) detected three cointegrating relationships, including the spread between the long and the short-term nominal interest rates, the long-term real interest rate, and the long-run money demand. Determinants driving M3 were weakly exogenous with respect to the long-run parameters. Long- and short-run parameter stability could not be rejected in the conditional single equation error-correction model for money demand. The cointegration vectors and the multivariate set-up were largely confirmed in the work by Brand and Cassola (2004). Deviations from the equilibrium mapped out by the three equations turned out to be relevant to explain the joint dynamics of inflation, income, money and interest rates.

At the same time, another model was developed by Calza, Gerdesmeier and Levy (2001) which was employed by the ECB as workhorse for assessing M3 developments since 2001. In this model, money demand was embedded in a cointegrated VAR model including M3 in real terms, real income and the spread between the short-run interest rate and the own rate of M3 holdings. The latter variable was calculated as the weighted average of the rates of return on individual M3 components, where weights are chosen according to the relative importance of each sub-component in M3.

A few years later, evidence emerged that extending the sample to include years after 2004 would cause instability in the existing money demand models, as the cointegration properties between money and the fundamentals tended to be lost. This led a first group of authors to look more closely at relationships between the core components instead of the original variables, the former generated by filter techniques (see Gerlach (2004) and Neumann and Greiber (2004)). At the same time, other studies included additional variables in the long-run relationship, among which were measures of financial volatility (Carstensen, 2006). In the same vein, Dreger and

Wolters (2010) provided empirical evidence that cointegration could be restored if inflation was allowed to enter the opportunity costs. Moreover, since portfolio shifts between money and other assets might have affected the long-run parameters over the last decade, financial wealth became increasingly important to explain the evolution of real money balances, as shown in Greiber and Setzer (2007) and Beyer (2009). According to Dreger and Wolters (2009), real house prices are especially suited to proxy the wealth effect, but their inclusion is tightly linked to an upward shift in the income elasticity in 2002. It coincides with the introduction of the euro to the public and could have signalled a rise in permanent income, caused by deeper monetary integration. Finally, De Santis, Favero and Roffia (2013) re-stablished a stable money demand relationship in the context of a portfolio-balance approach, where investors hold a diversified portfolio with money, domestic and foreign stock and long-term bonds.

All these studies show that money demand models which include variables in addition to the classical ones can restore stability in the money demand function. However, they might be less suited from the policy perspective, as they are more difficult to apply. In addition, models such as those based on core components may be of limited value for monetary analysis, given the well-known filtering problems at the end of the sample. Furthermore, improvements have been introduced *ex post* over the years, i.e. after detecting deficits of the standard specifications. In this respect, the inclusion of additional variables can indeed provide a better empirical fit, but this does not necessarily imply that the extensions will remain valid on an *ex ante* basis. For example, while the increase in real house prices may have contributed to the strong monetary acceleration before the economic crisis, their subsequent fall could provide problems when explaining money demand. Therefore, this study investigates the performance of the basic specifications, taken the crisis period into account. Apart from a few exceptions, such as Dreger and Wolters (2015), money demand in the crisis period has not been analysed so far.

3 Money demand specification

The standard model of money demand postulates a long-run relationship between real money, real income and nominal interest rates (Walsh, 2010). A slightly different variant is used in the analysis, i.e.

$$(1) \qquad (m-p)_t = \delta_0 + \delta_1 y_t + \delta_2 sy_t + \delta_3 R_t + \delta_4 own_t$$

where *m* denotes nominal money balances, *p* is the price level, and *y* is real income. Opportunity costs of holding money are proxied by the nominal long-term term interest rate (*R*), while the own rate of money holdings (*own*) is also included, as several components of M3 are interest-bearing. A distinct feature is the inclusion of the *sy* term, where *y* is multiplied with a step dummy *s* equal to 1 after 2002 Q1 and 0 before. The variable accounts for the break in the income elasticity after the euro was introduced as the common currency to the public (Dreger and Wolters, 2009). All variables are measured in logs, except for the interest rates, which are measured in percentages per annum.

From a theoretical point of view, income should exert a positive effect on real money balances, as it approximates transaction and precautionary savings motives. The parameter of the *sy* variable captures the potential change in the income elasticity after the euro introduction. If the income elasticity has risen, probably as a response to higher permanent income perspectives due to the common currency area, the coefficient will be positive. The semi-elasticity of money demand with respect to the long-term interest rate is negative, as stressed by the liquidity preference theory. In periods of rising interest rates, money holdings become less attractive. In contrast, the semi-elasticity with respect to the own yield of money holdings is expected to be positive.

As the euro area was established in 1999, artificial data are used for the period preceding this date. In particular, euro area series are constructed through the aggregation of national time series, see Calza, Gerdesmeier and Levy (2001), Brand, Gerdesmeier and Roffia (2002), and De Santis, Favero and Roffia (2013). In principle, different aggregation methods can have an impact on the results. As stressed by Bosker (2006) the aggregation bias is substantial prior to 1983, but almost negligible thereafter. In addition, the European Monetary System started working in 1983, and financial markets have become much more integrated since then. Thus, the sample period considered in the analysis is 1983 Q1-2015 Q1, where quarterly seasonally adjusted data are used (for the data sources, see Favero, De Santis and Roffia, 2013). Figure 1 displays the variables' developments over the sample.

-Figures 1 and 2 about here-

To derive real money balances and real income, the respective nominal series are deflated by the GDP deflator (set equal to 1 in 2010 Q1). Long-term interest rates refer to government bonds with 10 years to maturity. Following Bruggemann, Donati and Warne (2003) the own rate of M3 is constructed as a weighted average of the national own rates, the latter calculated as a weighting average of the rates of return of the individual components included in M3.[3] The euro area aggregate series are calculated from the national series using M3 weights.

It should be noted that the own rate of money holdings is not identical to the short-term interest rate. This point is quite important as the latter has been often used in studies on euro area money demand to proxy for the own rate effect. Differences occur mainly in the first years of the sample period (Figure 2). Afterwards, rising financial integration led to a faster convergence of the rates of return. At the same time, the share of marketable components in the overall M3 aggregate has increased in many euro area member states.

[3] The rate of return of currency in circulation is assumed to be nil. For the rates of return on the various categories of deposits, such as overnight deposits, short time savings and time deposits, the ECB constructed a new set of harmonised interest rates for monetary financial institutions. The rate of return of marketable components is proxied by the three-month money market rate. The weights of the individual components reflect their share in M3.

4 Money demand in the euro area

4.1 *Cointegrated VAR evidence*

While the variables in the money demand relationship are non-stationary in levels, their first differences are stationary. This result has been established in many studies, and can be also inferred from the inspection of Figure 1. Therefore, a cointegration analysis is the appropriate way to proceed. Based on the cointegrated VAR model, the Johansen (1992) trace statistic indicates a cointegration rank of 1 (see Table 1), implying that the cointegrating vector is unique. The cointegrating vector is shown in Table 2, together with the feedback parameters that reveal the short-run response of the variables to temporary deviations from the long run. While the upper part of Table 2 shows the reduced rank estimates from the unrestricted system, the lower part holds the results after restrictions are embedded.[4]

-Tables 1 and 2 about here-

The cointegrating vector is in line with a standard money demand function. The point estimate of the income elasticity of 0.61 in the unrestricted model version is below unity, but according to the standard errors, not significantly different from the value implied by the quantity theory of money (i.e. 1.0). The coefficient of *sy* implies that the elasticity has slightly increased after the introduction of the euro to the public. In addition, the interest rates bear the correct sign, and the semi-elasticities appear to be equal in absolute value. As indicated by the feedback parameters, real money balances respond to deviations from the long run equilibrium. The coefficient is negative, as it is expected from a stable model.

[4] It should be noted that these results are robust. Recursive estimation does not reveal any problems neither with the cointegration rank, nor the cointegrating vector nor the feedback mechanisms. More detailed results are available from the authors upon request.

The unrestricted findings suggest several restrictions that should be implemented and examined by means of likelihood ratio tests. First, it is investigated whether the interest rate coefficients are absolutely equal and of opposite sign. The corresponding null hypothesis cannot be rejected (Chi-square(1)=0.136, *p*-value =0.712). Second, it is tested whether all variables are weakly exogenous with respect to the cointegrating relationship, except for real money balances. Again, this hypothesis is in line with the data (Chi-square(5)=4.776, *p*-value=0.444). Because money is the only variable that may restore the long-run equilibrium after the emergence of shocks, the cointegration vector can be fairly interpreted as a money demand function. Apart from the constant, the error correction term can be written as follows:

$$(2) \qquad ECM_{ML,t} = \left(m - p\right)_t - 0.850 y_t - 0.021 sy_t + 0.094\left(R - own\right)_t$$

where *ML* denotes that the parameters are estimated by maximum likelihood. Note that the income elasticity increases slightly if the restrictions are applied. While the point elasticity is quite appealing, it is lower than in almost all previous studies. For example, Dreger and Wolters (2015) report an income elasticity significantly above unity. This result is likely driven by measurement error, since the own rate is proxied by the short-term interest rate in their study.[5] Therefore, it can be concluded that the use of a carefully constructed own rate of money holdings can help to resolving the long standing puzzle of an income elasticity exceeding 1.0.

4.2 *Conditional error correction modelling*

From the perspective of policy-makers, a cointegrated VAR model might be unattractive, especially in real time. Furthermore, the cointegration vector is unique, and real money balances is the only variable that is not weakly exogenous. Under these circumstances, a condi-

[5] In fact, if the cointegration analysis is done with the three-month money market rate instead of the own rate of M3 holdings, the income elasticity will increase to 1.8.

tional single equation error correction model might provide a reasonable alternative. It also allows further tests on the model specification.

According to Stock (1987), the single equation model can be estimated by OLS in one single step, providing the long-run parameters jointly with the parameters of the short-run dynamics. Initially, the first-period lags of the cointegrated variables are included. The short-run dynamics are approximated by the contemporaneous changes, and the first four lags of the first differences of all variables and a constant are included. In addition, other variables like inflation (i.e. the quarter-on-quarter change in the GDP deflator (*dp*)) and house prices (*w*) are allowed to enter the short-run dynamics, as suggested by the literature. As unconventional monetary policies may also affect real money balances in the short run, they are proxied by various measures, like innovations in the ECB balance sheet or the CISS indicator to capture systemic stress in the financial system.[6] The initial model is then simplified step-by-step, by eliminating the variables with the lowest and insignificant *t*-values (at 10% probability). The final specification of the conditional error correction model is shown in Table 3. In order to be consistent with the systems approach, the interest rate restriction is applied, which is highly supported by the data (Chi-square(1)=0.016, *p*-value= 0.899).

-Table 3 and Figure 3 about here-

The implied error correction term

(3) $$ECM_{OLS,t} = \left(m - p\right)_t - 0.903y_t - 0.022sy_t + 0.093\left(R - own\right)_t$$

appears to be very similar to the one obtained with the cointegrated VAR. The two ECMs move in close parallel, with a correlation coefficient of 0.91 (Figure 3). Positive values of the ECM indicate a monetary overhang, i.e. the actual money stock exceeds the level implied

6 See Hollo, Kremer and Lo Duca (2012) and Peersman (2014) for a discussion of these series.

by the money demand equation. At the end of the sample there are no signs of excess liquidity in the euro area, although the unconventional monetary policy measures have led to a substantial increase in the monetary base. However, these impulses have not been transferred to the M3 aggregate so far, as the monetary transmission process seems to be still impaired. This conclusion is reinforced by the fact that the proxies for unconventional monetary policies do not play any role for the evolution of real money balances.[7]

The specification tests do not point to any shortcomings (Table 3). The residuals are well-behaved, as they are normal, homoscedastic and not autocorrelated. The parameters are stable, as indicated by the Chow forecasting tests and the Cusum of Squares (Figure 4). Taking all this information together, we can conclude that a standard money demand function can be established for the euro area. The long-run coefficients bear the correct sign and are of reasonable size. The findings hold even if the most recent period is taken into account.

4.3 *Testing the relevance of the Lucas critique*

Notwithstanding the results mentioned above, the model might be of limited value for policy use, in case the Lucas critique applies. In the latter case, the parameters might vary with the policy regime. From an empirical point of view, however, the critique is not relevant if the explanatory variables under consideration can be treated as superexogenous. In that case, the conditional model parameters are expected to remain constant even if the parameters of the marginal processes describing the weakly exogeneous regressors are affected by policy shifts (Favero and Hendry, 1992).

-Table 4 about here-

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[7] As variables for unconventional monetary policy measures are not available before 1999, their power to explain M3 money demand has been investigated in a shorter sample period.

Testing for superexogeneity can be carried out via a two-step approach. First, marginal processes need to be specified for the weakly exogenous variables. As the own rate, the long-term interest rate and the inflation rate enter the single equation error correction model contemporaneously, their marginal processes are those considered for the test (Table 4). After adding impulse dummies for potential policy shifts and other shocks, the equations are well behaved. The residuals do not exhibit autocorrelation or heteroscedasticity. The parameters are also constant over time.

Second, the conditional model is extended with the explanatory variables from the marginal processes (Ericsson and Irons, 1995). In this case, if the additional variables do not affect the conditional model, the null hypothesis of superexogeneity cannot be rejected for the variables under study. The test results in *F*-values for the Wald test of 0.39 (*p*-value 0.97) for the own rate of M3 holdings, 0.67 (*p*-value 0.61) for the long-term interest rate and 0.85 (*p*-value 0.62) for the inflation rate.[8] This suggests that the money demand equation is not influenced by the Lucas critique and can provide an instrument for policy analysis.

5 Forecasting inflation with monetary aggregates

Monetary aggregates can provide information on future inflation if they are shown to improve inflation forecasts when their information content is taken into account. In this context, we test the predictive ability of monetary aggregates with respect to the consumer price (*pc*) inflation. Subsequently, we test the accuracy of such relationship over various forecasting horizon ranging from one to two and three years (see Dreger and Wolters (2014)). The inflation rates are defined as follows:

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[8] It could be argued that this result emerges since the null hypothesis states that the marginal regressors are jointly equal to 0 in the conditional model. Especially in the marginal process governing the development of inflation, many impulse dummies are needed to obtain a stable model. In principle, a few of them may be significant in the conditional model. However, the *t*-values of the marginal regressors are all well below the critical values.

(4) $$\pi_t^k = 4\log\left(pc_t / pc_{t-k}\right)/k \quad , \quad k = 1,4,8,12 \text{ quarters}$$

Out-of sample forecasts refer to the annual change of the consumer price index, i.e. (pc/pc_{t-4}), as well as average annual inflation rates over the two- and three-year horizon (pc/pc_{t-8}, pc/pc_{t-12}). These measures can be considered to be of high relevance for the monetary authorities, as they reveal information on the inflation potential over the medium and long run. Short-run changes in high volatile prices are removed if multiannual measures are considered. In order to mimic the situation faced by the policy-makers when forecasting inflation, the following direct approach is selected:

(5) $$\pi_{t+k}^k = \alpha(L)\pi_t^1 + \beta x_t + u_{t+k} \quad , \quad k = 4,8,12 \text{ quarters}$$

where $\alpha(L)$ is a lag polynomial, ensuring that the equations are balanced. Future inflation for k=4, 8 or 12 quarters ahead is then predicted by current and lagged quarterly inflation up to order 3 and variables known at the time the forecast is made (x). Lagged values of x do not contribute significantly in the forecasting equation and are thus excluded. Since the forecast error u follows a moving average process of order k-1, the autocorrelation and heteroscedasticity consistent covariance estimator proposed by Newey and West (1987) is used to evaluate the significance of the regression parameters.

The forecasting performance of this model is assessed via-à-vis a benchmark which is represented by a specification in which future inflation is predicted on the basis of only current and lagged inflation. Alternative models are obtained by adding other variables to those contained in the benchmark model. In particular, the first alternative model is based on annual M3 growth, while the second one includes the error correction term (i.e. the excess liquidity) at time t. Here, the ML variant is considered, although the OLS alternative leads to very similar results. To examine whether unconventional monetary policy measures can improve the forecasting performance, the error correction specification is extended by the annual change of the ECB balance sheet (see also Curdia and Woodford (2011) for similar considerations).

The forecasting performance is evaluated in an out-of-sample exercise, which mimics the actual situation faced by the forecasters. Due to the stability of the long-run money demand equation, differences in the error correction terms between the full sample and the corresponding subsamples can be largely neglected. The forecasts are obtained in a recursive manner. The first estimation subsample is 1983 Q1-2008 Q4 and the forecast subsample is 2009 Q4-2015 Q1 if annual inflation rates are predicted. After producing the forecast for 2009 Q4, the estimation period is extended by one quarter (1983 Q1-2009 Q1) and the forecast for 2010 Q1 is made. This process is repeated until the end of the sample (the last forecast subsample is 2014 Q1-2015Q1). Overall, 26 annual, 22 biennial and 18 triennial forecasts are derived.

The forecast accuracy is evaluated by the Theil-*U* coefficient, i.e. the root mean square forecast error of a specific method divided by its counterpart from the benchmark model (Table 5). For robustness, Theil's *U* is also expressed for mean absolute forecast errors. In general, the average root mean square forecast error exceeds the mean absolute forecast error due to possible outliers. Theil ratios below (above) unity indicate an improvement (worsening) relative to the benchmark. To assess the significance of the difference in the root mean square forecast or mean absolute forecast errors, the *p*-values from the Diebold-Mariano (1995) test are reported. This one-sided test explores the null hypothesis that rival forecasting models have equal predictive accuracy against the alternative that a particular method outperforms the benchmark. Simulation evidence indicates that the Diebold-Mariano statistic can be compared to critical values from the standard normal distribution, when forecasts are generated under rolling or recursive schemes (Giacomini and White, 2006). The Diebold-Mariano statistic is adjusted for small samples (see Harvey, Leybourne and Newbold (1997)).

-Table 5 about here-

The predictive accuracy can be significantly improved for all horizons if the forecasting equation is enriched by including the excess liquidity measure. In particular, significant gains are visible at the longer horizons. This means that fundamental information becomes gradually more important if the forecasting horizon increases. The respective equation leads to an average root mean square forecast error which is about 60 percent below the benchmark at the three-year horizon. Although the benefits might be exaggerated due to the specific crisis experience, these results underpin the usefulness of excess liquidity measures. It is also worth noting that the predictive accuracy does not increase if the forecasting equation is extended with the inclusion of the annual M3 growth. Hence, money *per se* does not seem to improve inflation forecasts, while it does when excess liquidity measures are considered. Furthermore, if the excess liquidity model is further extended with the inclusion of unconventional monetary policies, no additional gains are obtained. This also implies that these measures have been rather irrelevant for inflation so far, although this result is based on a backward-looking assessment.

The fact that the ECB's unconventional monetary policy impulses are reflected in strong increases in M0 rather than in comparable developments of M3 seems to be confusing at first glance, since the ECB has explicitly announced them with the aim of spurring credit growth and bringing inflationary developments more in line with the ECB's definition of price stability. At the same time, date seem to show that the banking system (facing a process of deleveraging of balance sheets) has often used the additional cheap liquidity (at least in the past) to repair the balance sheets by compensating on losses on subprime and other bad loans rather than providing new loans (see ECB (2015) for a more detailed analysis).

6 Conclusions

This paper investigates the stability of the demand for M3 in the euro area over the past three decades, including the recent economic crisis. The cointegration analysis leads to a well-identified model, which includes real money balances, income, the long-term interest rate and the own rate of M3 holdings. In addition, a one-time permanent break in the income elasticity is allowed, as the introduction of the euro to the public might have raised the permanent income perspectives. The specification appears to be robust against the Lucas critique of a policy-dependent parameter regime. In contrast to the common view, money demand does not show signs of instability in periods when interest rates reach the zero lower bound. Deviations of money demand from its equilibrium level do not indicate substantial inflation pressure at the end of the sample. Finally, excess liquidity models are able to outperform the autoregressive benchmark, as they deliver more accurate consumer inflation forecasts, especially at the longer horizons. The inclusion of unconventional monetary policy measures does not change these findings.

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Figure 1: Variables in the analysis

Levels

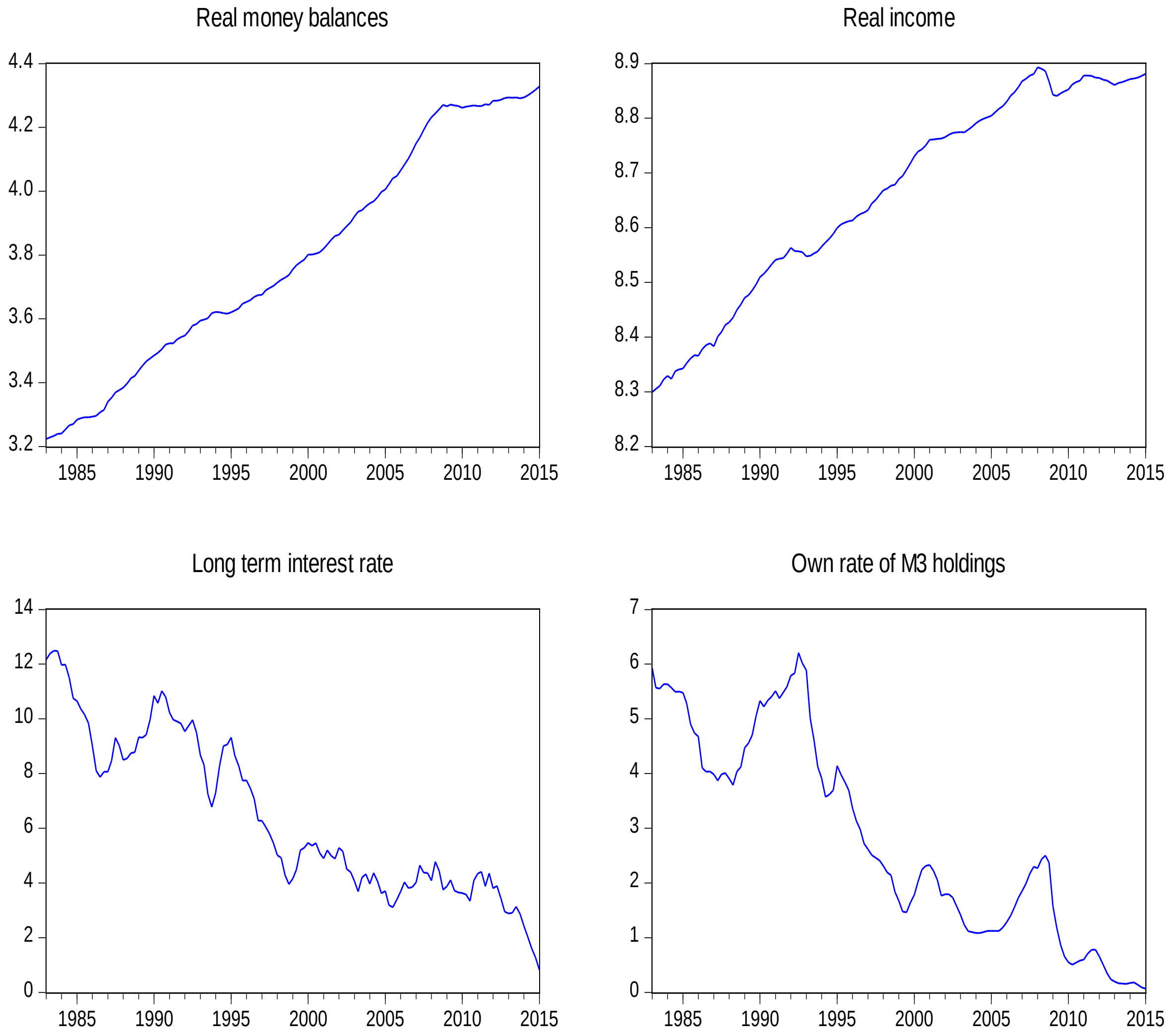
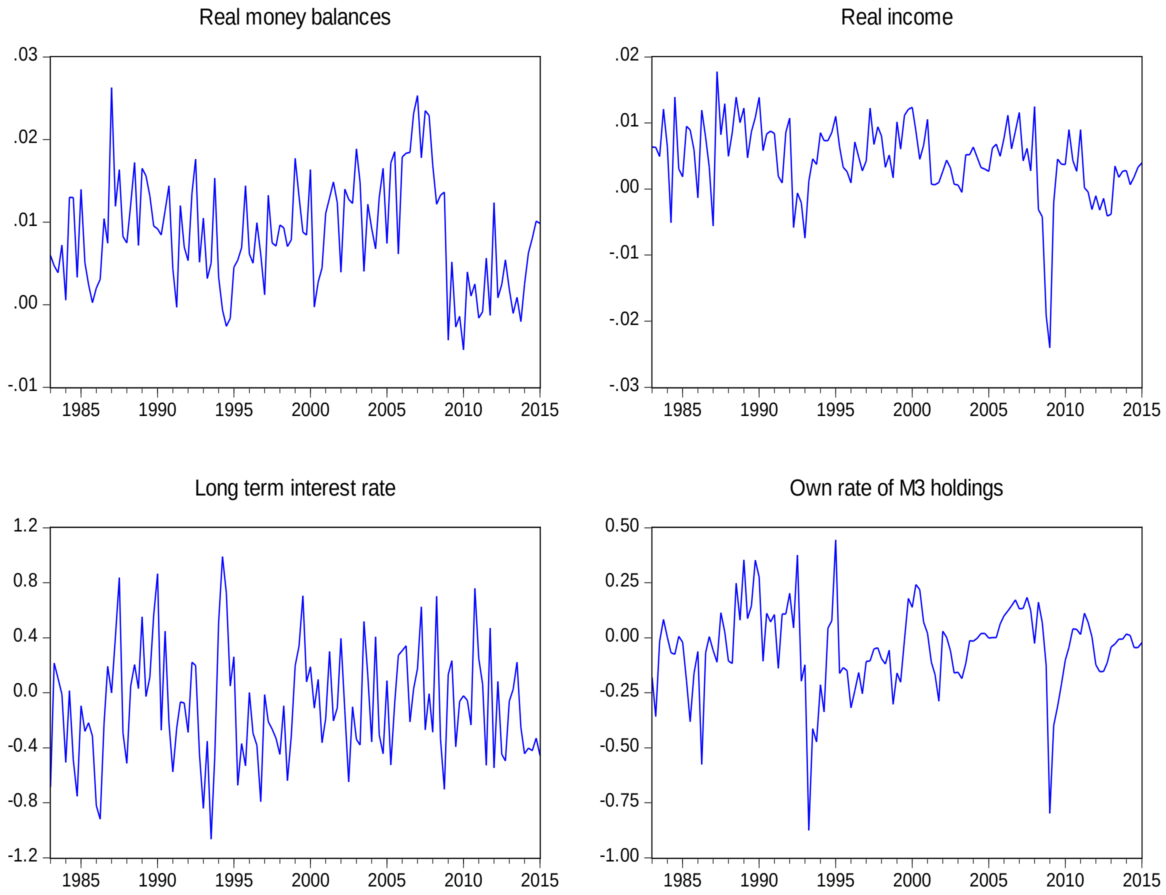


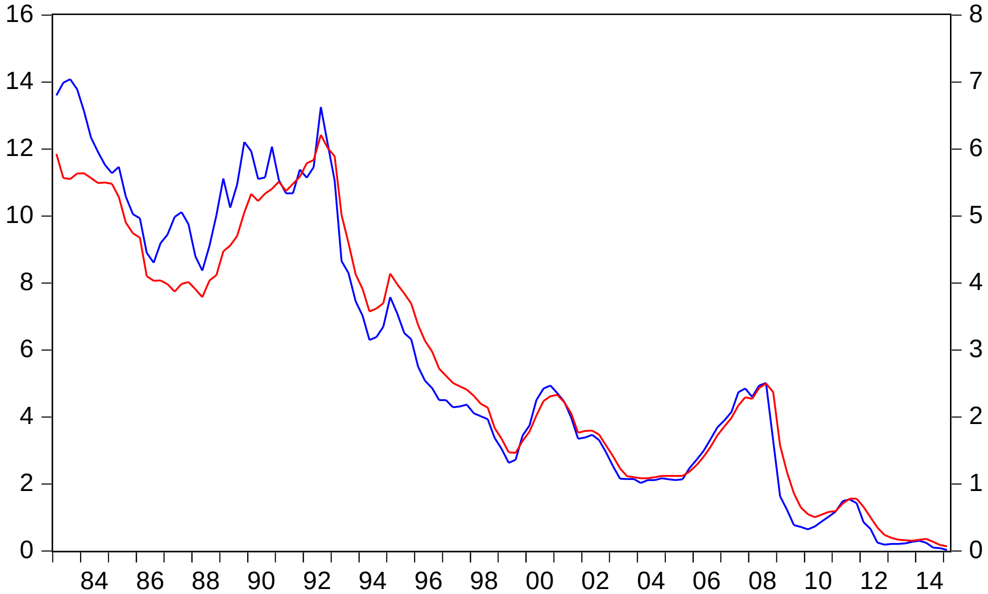
Figure 1 (cont'd)

First differences



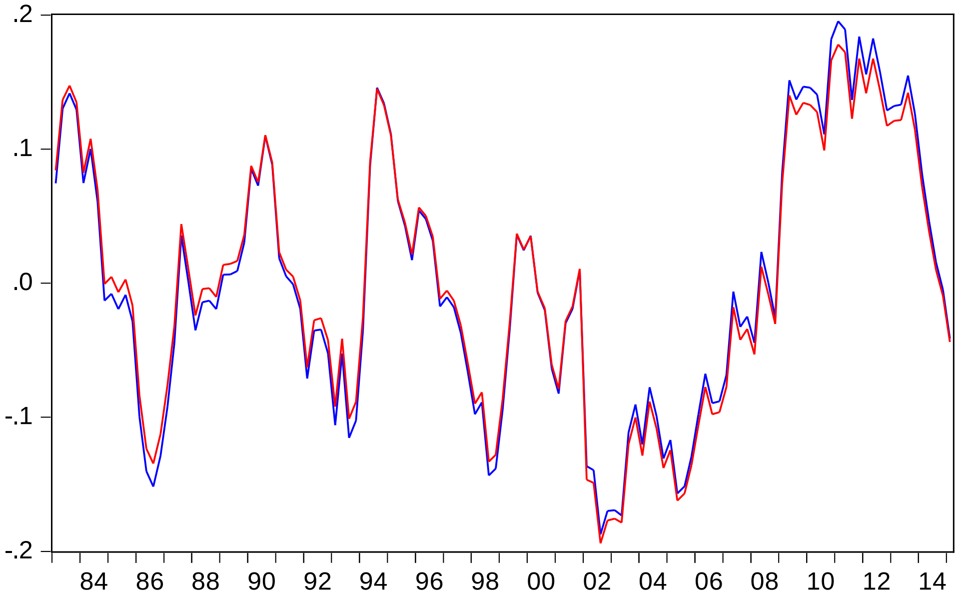
Notes: Sample period 1983 Q1-2015 Q1. Real money and real GDP in logarithms and deflated by the GDP deflator.

Figure 2: Three-month money market rate (blue) and own rate for M3 holdings



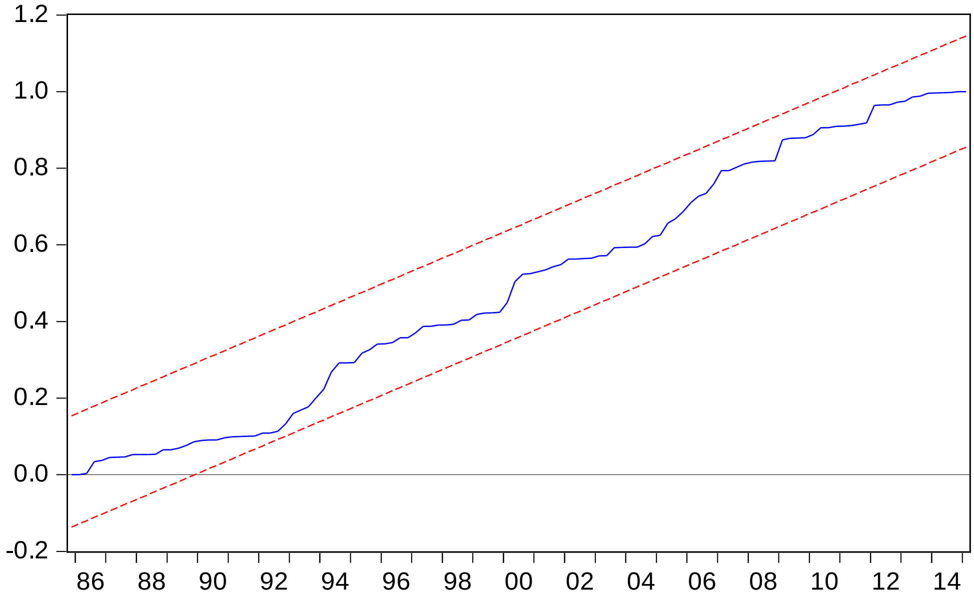
Note: Sample period 1983 Q1-2015 Q1.

Figure 3: Excess liquidity in the euro area



Notes: Sample period 1983 Q1-2015 Q1. Mean adjusted error correction terms obtained by ML (blue) and OLS methods.

Figure 4: Cusums of squares of the error correction model



Notes: Sample period 1983 Q1-2015 Q1. Dashed lines represent 0.05 significance levels.

Table 1 Cointegration rank in the money demand model

$r \leq 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$	$r \leq 4$
79.53 (0.006)	40.72 (0.199)	14.91 (0.789)	7.901 (0277)	2.838 (0.092)

Notes: Sample period 1983 Q1-2015 Q1. The model is based on real money balances, real income, long-term interest rate and the own rate for M3 holdings. Trace statistics for the null hypothesis that cointegration rank *r* is less than or equal *p*, corrected for small sample bias (Johansen, 2002) Models include unrestricted constant. Lag order of the underlying VAR model in levels is equal to 2 (AIC), *p*-values in parantheses.

Table 2 Cointegration vector and feedback parameters

A Unrestricted model

Cointegration vector		Feedback parameters	
m-p	1.000	Δ(*m-p*)	-0.024 (0.005)
y	-0.613 (0.241)	Δ*y*	-0.001 (0.005)
*y**	-0.023 (0.006)	Δ*y**	0.747 (0.713)
R	0.142 (0.026)	Δ*R*	-0.261 (0.353)
own	-0.155 (0.034)	Δ*r*	0.235 (0.136)

B Restricted model

Cointegration vector		Feedback parameters	
m-p	1.000	Δ(*m-p*)	-0.036 (0.006)
y	-0.850 (0.177)	Δ*y*	0
*y**	-0.021 (0.005)	Δ*y**	0
R	0.094 (0.021)	Δ*R*	0
own	-0.094 (0.021)	Δ*r*	0

Notes: Sample period 1983 Q1-2015 Q1. Standard errors are in parentheses.

Table 3 Error correction model for M3 money demand

Regression coefficients		Specification and stability tests			
Constant	-0.088 (0.082)	R2	0.570	CF(07Q1)	0.816 (0.740)
$\Delta(m\text{-}p)_{t-1}$	0.241 (0.079)	LM(1)	2.301 (0.132)	CF(08Q1)	0.714 (0.848)
$\Delta(m\text{-}p)_{t-2}$	0.213 (0.077)	LM(4)	1.537 (0.196)	CF(09Q1)	0.819 (0.709)
ΔR_t	-0.003 (0.001)			CF(10Q1)	0.633 (0.884)
Δdp_t	-0.002 (0.001)	ARCH(1)	0.520 (0.472)	CF(11Q1)	0.593 (0.891)
Δdp_{t-1}	-0.002 (0.001)	ARCH(4)	0.738 (0.568)	CF(12Q1)	0.719 (0.741)
Δown_t	0.013 (0.002)			CF(13Q1)	0.345 (0.958)
$(m\text{-}p)_{t-1}$	-0.025 (0.007)	JB	1.330 (0.514)	CF(14Q1)	0.090 (0.994)
y_{t-1}	0.022 (0.012)	RESET(1)	1.412 (0.161)	CF(15Q1)	0.241 (0.810)
sy_{t-1}	0.001 (0.0002)	RESET(2)	1.792 (0.171)		
$(R\text{-}own)_{t-1}$	-0.002 (0.001)	RESET(3)	1.259 (0.292)		

Notes: Sample period 1983 Q1-2015 Q1. Standard errors of the regression coefficients are in parentheses. R2 adjusted R-squared, JB=Jarque-Bera test for normality of residuals, LM=Lagrange multiplier test for no autocorrelation in the residuals up to specified order, ARCH=ARCH test for no conditional heteroscedasticity up to specified order, RESET=Ramsey specification test, CF=Chow forecast test for no structural break at specified quarter, *p*-values for specification and stability tests in parantheses.

Table 4: Marginal processes for weakly exogeneous variables

Dependent: Δown_t

Con	Δown_{t-1}	Δown_{t-2}	ΔR_{t-1}	Δdp_{t-5}	Δy_{t-1}	$i862$	$i894$	$i902$
-0.056 (0.014)	0.215 (0.064)	0.199 (0.058)	0.131 (0.026)	0.016 (0.007)	0.118 (0.018)	0.118 (0.018)	0.271 (0.107)	-0.442 (0.111)

$i923$	$i932$	$i942$	$i951$	$i952$
0.388 (0.109)	-0.533 (0.110)	-0.353 (0.111)	0.382 (0.106)	-0.400 (0.109)

R2=0.713

ARCH(1)=0.13 (0.72)	ARCH(1)=0.28 (0.89)	LM(1)=0.54 (0.47)	LM(4)=0.39 (0.82)

Dependent: ΔR_t

Con	ΔR_{t-1}	Δown_{t-4}	Δdp_{t-2}	Δdp_{t-4}
-0.087 (0.035)	0.267 (0.083)	-0.353 (0.169)	0.056 (0.024)	0.044 (0.024)

R2=0.121

ARCH(1)=0.15 (0.70)	ARCH(4)=1.48 (0.21)	LM(1)=1.22 (0.27)	LM(4)=0.51 (0.73)

Dependent: Δdp_t

Con	Δdp_{t-1}	Δdp_{t-2}	ΔR_{t-5}	Δy_{t-1}	$i842$	$i861$	$i862$	$i863$
-0.304 (0.086)	-0.764 (0.069)	-0.342 (0.064)	0.380 (0.172)	0.618 (0.115)	-2.868 (0.741)	1.767 (0.753)	-2.142 (0.737)	-2.047 (0.766)

$i871$	$i872$	$i874$	$i883$	$i904$	$i912$	$i931$	$i934$
-5.830 (0.733)	2.051 (0.865)	2.908 (0.798)	-1.931 (0.737)	-1.806 (0.734)	2.681 (0.763)	2.504 (0.742)	-1.805 (0.749)

$R2=0.737$

ARCH(1)=0.87 (0.34)	ARCH(4)=1.39 (0.24)	LM(1)=0.64 (0.43)	LM(4)=0.88 (0.48)

Notes: Sample period 1983 Q1-2015 Q1. R2=R squared adjusted, LM(*k*)=Lagrange multiplier test for no autocorrelation in the residuals up to order *k*, ARCH(*k*)= LM test for no conditional heteroscedasticity up to order *k*. In parentheses standard errors for regression coefficients, *p*-values for test statistics. Impulse dummies *i*zzq are equal to 1 in year zz and quarter q and 0 otherwise.

Table 5: Out-of-sample forecasting performance of different models

A Root mean squared forecast error

Horizon	Benchmark	Money growth	Excess liquidity	Excess liquidity & UMP
4	1.70	1.01 | 0.55	0.83 | 0.21	0.90 | 0.36
8	1.77	0.99 | 0.44	0.58 | 0.01	0.51 | 0.01
12	1.63	0.96 | 0.28	0.36 | 0.00	0.28 | 0.00

B Mean absolute forecast error

Horizon	Benchmark	Money growth	Excess liquidity	Excess liquidity & UMP
4	1.24	1.12 | 0.85	0.96 | 0.42	1.05 | 0.57
8	1.51	1.02 | 0.58	0.59 | 0.00	0.47 | 0.01
12	1.40	0.96 | 0.32	0.37 | 0.01	0.25 | 0.00

Notes: UMP=Unconventional monetary policy, excess liquidity according to the ML estimates. Root mean squared forecast error (RMSFE) and mean absolute forecast error (MAFE) for different forecast horizons refer to the autoregressive benchmark, expressed in percent. The three columns on the right report the RMSFE or MAFE relative to that of the benchmark. The entries on the left-hand side denote the Theil-*U* coefficients, the right-hand figures denote the *p*-values of the Diebold-Mariano (1995) test statistic on equal predictive accuracy. The Harvey, Leybourne and Newbold (1997) correction for small samples is applied.