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*The analyses, opinions and findings of these papers represent the views of the authors, they are not necessarily those of the Banco de Portugal.*

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# Is the euro area M3 abandoning us?\*

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## Abstract

This paper reassesses the role of the M3 aggregate for monetary policy purposes in the euro area. Using data until 2006Q4 it is shown that the M3 aggregate ceased to display the empirical properties that supported its prominent role in the ECB's monetary policy strategy. On the one hand, when the most recent data are used in the analysis there is strong evidence of cointegration breakdown in the M3 money demand models as well as in the "two-pillar Phillips curves" with filtered data. On the other hand, the leading indicator properties of M3 for inflation in the area have also deteriorated markedly in the most recent years. This is supported by evidence both in the time and frequency domains.

JEL: E3; E4; E5.

Keywords: M3; Euro area; Cointegration breakdown; Leading indicator properties; Frequency domain.

## 1 Introduction

The choice of M3 as the prominent monetary aggregate in the ECB's monetary policy framework was based on two criteria (see ECB, 2004a, Coenen and Vega, 2001, and Trecroci and Vega, 2000). First, this aggregate exhibited a remarkably stable long-run relationship with its traditional determinants. Second, it displayed leading indicator properties regarding future inflation in the medium term. These two criteria are well grounded on the theoretical literature and aim to ensure that money can be used as a reliable information variable (see Friedman, 1970 and Issing et al.,2001). It is thus not

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surprising that these criteria were reaffirmed in the ECB's evaluation of the monetary policy strategy in 2003 (ECB, 2003).

Since the inception of EMU, numerous papers were produced supporting the idea that M3 exhibited the two above criteria<sup>1</sup>. However, the monetary dynamics in the euro area since 2001 – with M3 gradually accelerating to rates of growth above 10 per cent, in a context of broadly stable inflation around 2 per cent, anchored inflation expectations and moderate rates of economic growth – has raised questions as to whether the most recent evidence continues to support that conclusion. In fact, several recent contributions have raised doubts concerning both the stability of the long-run M3 money demand relation and the leading indicator properties of M3, when the most recent data are included in the analysis<sup>2</sup>. More recently, it has been claimed that the M3 aggregate could still be useful provided the long-run fluctuations of inflation were closely related to long-run fluctuations in M3 growth (see Assenmacher-Wesche and Gerlach, 2006a).

In this paper, we aim to re-assess this debate on the usefulness of M3 as an indicator of risks to price stability in the euro area. In particular, using a dataset updated until 2006Q4, we test the stability of several long-run M3 money demand models, the stability of the relation between the low frequencies of M3 growth and inflation and, finally, the existence of leading indicator properties of M3 for inflation, both in the time and frequency domains.

With respect to the long-run M3 money demand three models were re-evaluated: two versions of the model suggested in Calza, Gerdesmeier and Levy (2001) [hereafter CGL, 2001], which has been extensively used by the ECB for monetary policy analysis, and the one suggested in Carstensen (2004, 2006) which attempts to endogenise the so-called portfolio shifts<sup>3</sup>. To anticipate the results, we show that these money demand

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<sup>1</sup>Studies aiming at uncovering a stable long-run M3 money demand equation in the euro area include Coenen and Vega (2001), Brand and Cassola (2000), Calza, Gerdesmeier and Levy (2001), Cassola and Morana (2002), Bruggeman, Donati and Warne (2003), Carstensen (2004, 2006), Dreger and Wolters (2006), Landesberger (2007) and Greiber and Setzer (2007). Studies aiming at establishing the leading indicator property of M3 for inflation include Trecroci and Vega (2000), Altimari (2001) and Fisher et al. (2006).

<sup>2</sup>The stability of the long-run M3 money demand relation has been subject to critique in Alves et al., 2006 and Bordes et al., 2007. The gradual fading of the leading indicator properties of M3 growth with respect to future inflation when the most recent vintages of data are incorporated in the models are reported in OECD (2007), Hoffman (2006) and Lenza (2006).

<sup>3</sup>The ECB's explanation for the acceleration of M3 in the period 2001-2003 relies on the idea that increased uncertainty in the stock market led to portfolio adjustments towards more liquid and safer assets included in M3 (the "portfolio shifts"). In this context, and as a complement to the official M3 aggregate, the ECB has built a new aggregate, the so-called "M3 corrected for the impact of portfolio shifts" [see ECB, 2004]. However, using such an aggregate for monetary analysis raises several important questions. First, the correction is completely ad-hoc, based on simple non-causal time series models, implying that in fact we do not know what the money stock would have been in the absence of such

models exhibit strong signs of instability or cointegration breakdown when data until the end of 2006 are used in the analysis. The emergence of cointegration breakdown has the implication that a stable long-run relation linking M3, prices and the level of activity, as specified in these models, ceased to exist. Cointegration breakdown also implies that the monetary indicators based on the residuals of the cointegrating regressions, such as the monetary overhang/shortfall and the real money gap, lose their interpretation as excess liquidity indicators.

In order to investigate whether long term (or low frequency) fluctuations in inflation are closely associated with long term (low frequency) movements in M3 growth we re-evaluate the approach recently suggested in Assenmacher-Wesche and Gerlach (2006a). In this case, end-of-sample stability tests show that stability is also at stake for these models when data until the end of 2006 are incorporated in the analysis.

In what concerns the leading indicator properties of M3, we update the out-of-sample forecasting exercise of Altimari (2001), based on the Stock and Watson (1999) methodology. We show that the forecasting performance of M3 has substantially deteriorated in recent years and that the recent dynamics of M3 worsen the inflation forecasting performance of a simple random-walk model. Further, the deterioration of the information content of M3 in the recent past is also confirmed in the frequency domain, using the test proposed in Breitung and Candelon (2006).

In sum, we show that M3 ceased to comply with the Issing et al. (2001) criteria that “the chosen aggregate must have a stable, predictable long-run relationship with prices, as well as good leading indicator properties in the medium term”.

The remainder of the paper is organised as follows. Section 2 re-evaluates several M3 money demand models for the euro area and discusses the implications for the excess liquidity indicators stemming from the cointegration breakdown. Section 3 revisits the Phillips-curve approach suggested in Assenmacher-Wesche and Gerlach (2006a). Section 4 assesses the leading indicator properties of M3 and discusses their robustness. Section 5 presents the main conclusions.

## **2 Stability of the long-run money demand function**

In this section we investigate whether the long-run money demand equation for M3 remains stable when data until 2006Q4 are added to the sample. For that purpose we

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portfolio shifts. Second, it is used under the assumption that the existing models (including the estimates of the parameters) would have remained valid after 2001/2002 in the absence of such shifts, something that cannot be investigated. Finally, as we will show, cointegration breakdown in the long run money demand, for the most recent period, cannot be explained by portfolio shifts.

specially focus on the CGL (2001) and Carstensen (2004, 2006) models because the first has been used by the ECB in monetary assessments and the second is an extension of the original CGL (2001) model that endogeneizes the so-called portfolio shifts that started to disrupt the official M3 aggregate after mid-2001.

Based on the money demand function developed by Calza, Gerdesmeier and Levy (2001), the ECB conveyed the message that long-run M3 demand has remained stable despite the strong growth rate displayed by that aggregate after mid-2001, well above the  $4\frac{1}{2}$  per cent reference value (ECB, 2004). In contrast, Carstensen (2004) conducted a thorough analysis on the existing models of money demand (including the original version of the CGL model) and concluded that they were generally stable when data until 2001 are considered but most of them exhibit instability problems when data after 2001 are added to the sample. As an alternative Carstensen (2004, 2006) suggests a new model which appears to be stable when estimated with data until the second quarter of 2003 (the maximum sample available to the author).

The original version of the CGL (2001) model is a VAR comprising real M3, real GDP and the opportunity cost of M3 (the spread between the short-term market interest rate and the own rate of M3), with two lags in the levels of the variables. A more recent version of the CGL (2001) model (ECB, 2004) includes in addition the following exogenous stationary variables affecting only the short-term dynamics: one quarter-lagged change in oil prices and in the yield curve (defined as the spread between the long-term government bond yield and the short-term market interest rate) and the first difference of the annualised quarterly inflation rate (based on the GDP deflator). In what follows these two models will be denoted the “original version” and the “revised version” of the CGL model.

The model suggested in Carstensen (2004, 2006) is an extension of the original version of the CGL (2001) model which, besides real M3, real GDP and the spread between the short-term market interest rate and the own rate of M3, also includes two stock market variables. These two additional variables are the stock market volatility and the spread between equity returns and the own rate of M3. Thus, in the case of the model suggested in Carstensen (2004, 2006) the long-run money demand function can be written as

$$(m - p)_t = \beta_0 + \beta_1 y_t + \beta_2 (r_t^s - r_t^o) + \beta_3 (r_t^e - r_t^o) + \beta_4 z_t \quad (1)$$

where  $(m - p)_t$  stands for the log of the real money stock,  $y_t$  for the log of real GDP,  $r_t^s$  for the short-term nominal interest rate,  $r_t^o$  for the nominal own rate of M3,  $r_t^e$  for the

nominal equity return and  $z_t$  for the log of stock market volatility<sup>4</sup>. In this model all the individual variables are assumed to be integrated of order one<sup>5</sup>. The original version of the CGL model obtains by setting  $\beta_3 = \beta_4 = 0$ .

For the re-evaluation that follows we use data until the fourth quarter of 2006. The analysis of the models is conducted at two different levels. We start by looking in subsection 2.1 at the cointegration tests in order to investigate whether cointegration holds when more recent data are added to the sample. In subsection 2.2 we proceed by formally testing for cointegration and stability breakdown using the tests recently suggested in Andrews and Kim (2003, 2006).

## 2.1 Johansen cointegrating tests

Table 2.1 displays the results of the Johansen cointegration tests for the null of zero cointegrating vectors against the alternative of (at least) one cointegrating vector (with p-values in brackets) in the revised version of the CGL model, as defined above. The sample starts in 1980Q3 (the first two observations are used to account for the two lags of the model) and the end-of-sample varies from 2002Q1 to 2006Q4. Table 2.1 reports the p-values using both the asymptotic distribution (columns 2 and 3) and the small sample corrected critical values (columns 4 and 5). Following the discussion in the literature that suggests that the conventional asymptotic trace and maximum eigenvalue tests are subject to size distortions in small samples, we focus on the small sample corrected critical values<sup>6</sup>.

The CGL (2001) model was developed under the assumption of a single cointegrating vector. Looking at the cointegration tests in Table 2.1 we see that cointegration is lost in 2003, as none of the tests including data for 2003Q2 and thereafter leads to a rejection of the null of zero cointegrating vectors for a 10% test. Moreover, the evidence against cointegration accumulates steadily over the remainder of the sample. Using the

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<sup>4</sup>The nominal equity returns and the stock market volatility were computed as in Carstensen (2004, 2006). More specifically, the nominal equity returns are constructed as the annualised three-year log differences of quarterly nominal stock prices as measured by the Dow Jones Euro Stoxx50. In turn, the stock market volatility is constructed as the two-year average of the conditional variance estimated from a Garch model with t-Student innovations applied to daily yields of the nominal stock price index. Data for the remaining variables were kindly provided by ECB staff.

<sup>5</sup>The use of stock market volatility as a nonstationary variable in (1) is usually criticised on the grounds that a proxy for uncertainty should be stationary (see, for instance, Dreger and Volters, 2006). Carstensen is aware of such criticism, but argues that the relevant issue is that the series behaves like a nonstationary variable in the given sample. In this paper we do not take a stand on this issue as our single purpose is to investigate whether the model suggested in Carstensen (2004, 2006) is robust to the extension of the sample until the end of 2006.

<sup>6</sup>The computations in this section were carried out using PcGive 10.

**Table 1:** Johansen cointegration tests for the CGL model (revised version)

	Trace test	Max test	Trace test	Max test	Weak exog. GDP
	[Prob, asymp]	[Prob, asymp]	[Prob, small sample]	[Prob, small sample]	[Prob, asymp]
80Q3-02Q4	30.3 [0.04]*	18.2 [0.12]	28.2 [0.08]	17.0 [0.18]	0.7 [0.41]
80Q3-03Q2	26.6 [0.12]	18.0 [0.14]	24.8 [0.17]	16.8 [0.19]	2.9 [0.09]
80Q3-03Q4	25.9 [0.13]	17.8 [0.14]	24.3 [0.20]	16.7 [0.20]	2.2 [0.06]
80Q3-04Q4	21.9 [0.31]	17.3 [0.17]	20.6 [0.39]	16.2 [0.22]	7.1 [0.01]*
80Q3-05Q4	20.3 [0.41]	16.6 [0.20]	19.1 [0.50]	15.6 [0.26]	8.4 [0.00]*
80Q3-06Q4	18.3 [0.55]	14.4 [0.35]	17.3 [0.63]	13.6 [0.42]	8.7 [0.00]*

Note: \* marks significance at 95% level.

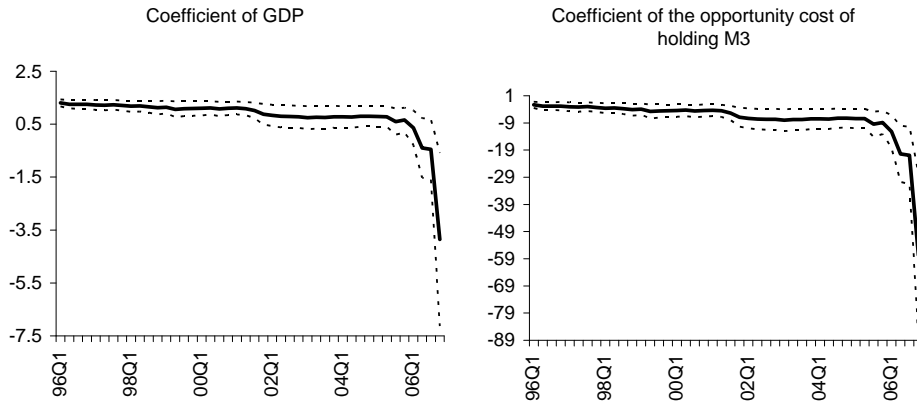
maximum sample period available we see that the null of zero cointegration vectors is not rejected even for a 40% test. More specifically, the p-values for the null of zero cointegrating vectors are 63% (trace test) and 42% (max test). Those figures are far beyond any acceptable level of significance used in the literature<sup>7</sup>.

Figure 1 depicts the recursive estimates of the long-run parameters associated with GDP and the opportunity cost with 95% confidence bands, obtained without re-estimating the short-run dynamics. Even though the simple inspection of recursive graphics does not constitute a formal stability test it nevertheless constitutes a very useful exercise as it allows a quick check of the evolution over time of the parameter estimates. By looking at Figure 1 it can be seen that the recursive estimates change significantly as more recent data are included in the sample. For instance, the point estimate for the GDP elasticity is 1.30 when the model is estimated with data until 1996Q1, but this estimate drops continuously as more data are added to the sample and becomes strongly negative (-3.86) when the full sample is used. In turn, the opportunity cost semi-elasticity increases (in absolute terms) from -2.32 in 1996Q1 to -57.7 in 2006Q4. In contrast with ECB (2004), Figure 1 was obtained without imposing any weak-exogeneity restriction given that the test of weak-exogeneity in the last column of Table 1 suggests that such a restriction ceased to be valid in 2004<sup>8</sup>. If, despite not being valid, we compute the recursive estimates by imposing the weak exogeneity restriction as in ECB (2004) we observe, for example, that the coefficient of GDP decreases from 1.36 in 1996Q1 to 0.72

<sup>7</sup>For space reasons we do not report the full set of cointegration tests for the original version of the CGL model. However the results are similar to those obtained for the revised version. For the full sample the small-sample corrected p-values for the null of zero cointegrating vectors are 58% (trace test) and 47% (max test).

<sup>8</sup>Note also that this test is valid only under the assumption of cointegration which, according to Table 1, can no longer be sustained.





**Figure 1:** Recursive estimates of the long-run coefficients in the CGL model (revised version) with 95 per cent confidence bands.

in 2006Q4.

As regards the model suggested in Carstensen (2004, 2006) Table 2.1 again displays the results of the Johansen cointegration tests for the null of zero cointegrating vectors against the alternative of (at least) one cointegrating vector (with p-values in brackets). The sample used in the tests starts in 1980Q3 (the first two observations are used to account for the two lags of the model) and the end-of-sample varies from 2003Q2 to 2006Q4.

From Table 2.1 we conclude that the Carstensen specification does a good job, as far as cointegration is concerned, until the first half of 2005. However when data for the second half of 2005 and after are added to the analysis the model deteriorates and cointegration is lost. In particular, when the maximum sample period is used (data until 2006Q4) we clearly see that the null of zero cointegration vectors cannot be rejected as the p-values of the test are 61% (trace test) and 79% (max test).

Figure 2 displays the recursive estimates of the long-run parameters of the Carstensen's model, obtained without imposing any weak-exogeneity restriction<sup>9</sup>. As expected, the estimated coefficients start to exhibit some instability after the beginning of 2003, and such instability becomes especially significant by the end of the sample.

Intuitively we can understand the outcome of the cointegration tests of the CGL and Carstensen models by looking at Figures A1 to A6 in Annex A. From Figure A1 we see that the real money stock accelerates after 2001. As this acceleration is not accompanied by a significant acceleration of real GDP (Figure A3) or by a significant decline in the

<sup>9</sup>Imposing weak exogeneity of GDP has no significant implications on the estimated coefficients.

**Table 2:** Johansen cointegration tests for Carstensen's model

Sample	Trace test	Max test	Trace test	Max test
	[Prob, asymp]	[Prob, asymp]	[Prob, small samp.]	[Prob, small samp.]
80Q3-03Q2	77.9 [0.01]*	41.2 [0.00]*	69.5 [0.05]	36.7 [0.02]*
80Q3-03Q4	76.1 [0.01]*	37.6 [0.01]*	68.0 [0.07]	33.6 [0.05]
80Q3-04Q4	78.6 [0.01]*	36.4 [0.02]*	70.6 [0.04]*	32.7 [0.07]
80Q3-05Q2	78.4 [0.01]*	35.7 [0.03]*	70.6 [0.4]*	32.1 [0.08]
80Q3-05Q3	72.1 [0.03]*	29.8 [0.15]	65.0 [0.11]	26.8 [0.28]
80Q3-05Q4	70.1 [0.05]*	28.9 [0.18]	63.2 [0.15]	26.1 [0.33]
80Q3-06Q4	56.0 [0.38]	21.5 [0.65]	50.8 [0.61]	19.5 [0.79]

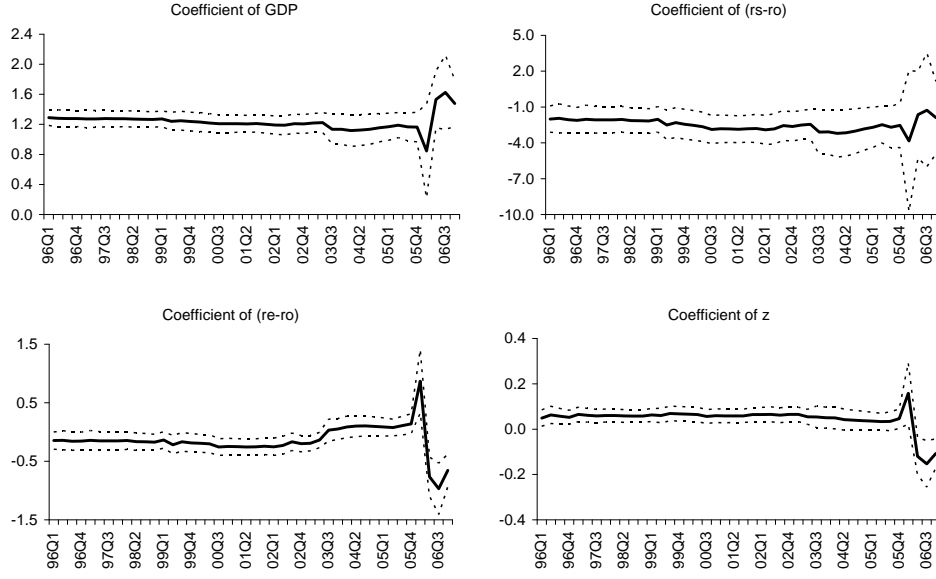
Note: \* marks significance at 95% level.

spread between the short term market rate and the own rate (Figure A4), the CGL model starts to perform poorer and poorer and eventually cointegration is lost in the first half of 2003, as Table 2.1 shows. On the other hand, Figures A5 and A6 show that the spread between the equity returns and the own rate,  $(r_t^e - r_t^o)$ , decreases and the volatility,  $z_t$ , increases until the beginning of 2003, which explains the good performance of the Carstensen specification in this period. However, after the first quarter of 2003 the spread  $(r_t^e - r_t^o)$  shows an increasing trend while volatility decreases. This, all else equal, should have brought about a decrease or at least a deceleration in money growth during this period which did not occur. This is why the model performs poorer after the first half of 2005.

## 2.2 Testing for stability and cointegration breakdown

In the previous subsection it was shown that when the most recent data are added to the sample, the evidence does not support the existence of cointegration in the CGL and Carstensen's models and the estimated long-run coefficients display significant changes. However, this analysis can be criticised on the grounds that cointegration tests may exhibit power problems and also that the recursive estimates of the long-run parameters with the corresponding 90 percent confidence intervals do not constitute formal stability tests. Thus, in this sub-section we address the issue in a more formal way by resorting to cointegration and stability breakdown tests recently suggested in the literature.

Andrews and Kim (2006) introduced two tests for cointegration breakdown that may occur at the end of the sample and thus are specially designed to investigate the problem at hand. The tests are conducted under the assumption that cointegration and stability of the long-run coefficients hold until a certain point in time and we want to investigate



**Figure 2:** Recursive estimates of the long-run coefficients in Carstensen's model with 95 per cent confidence bands.

whether there is a cointegration breakdown after that period. To fix ideas let us assume that the model is given by:

$$y_t = \begin{cases} x_t \beta_0 + u_t & \text{for } t = 1, \dots, T \\ x_t \beta_t + u_t & \text{for } t = T + 1, \dots, T + m \end{cases} \quad (2)$$

where  $x_t$  is the vector of the regressors and  $\beta$  the vector of the coefficients. The regressors are linear combinations of unit root random variables, stationary random variables and deterministic variables, such as a constant and a linear time trend.

Cointegration breakdown may occur due to a shift in the cointegration vector or to a shift in the errors from being stationary to being integrated of order one so that the null and alternative hypotheses of the tests conducted below are:

$$H_0 = \begin{cases} \beta_t = \beta_0 + u_t & \text{for all } t = T + 1, \dots, T + m \\ \{u_t : t = 1, \dots, T + m\} & \text{are stationary and ergodic} \end{cases} \quad (3)$$

$$H_1 = \begin{cases} \beta_t \neq \beta_0 & \text{for some } t = T + 1, \dots, T + m \text{ and/or} \\ \text{the distribution of } \{u_{T+1, \dots, T+m}\} & \text{differs from the distribution of } \{u_{1, \dots, m}\} \end{cases}$$

Under the null hypothesis the model is a well-specified cointegrating regression model for all  $t = 1, \dots, T + m$ . Under the alternative hypothesis the model is a well-specified

cointegrating regression model for all  $t = 1, \dots, T$ , but for  $t = T + 1, \dots, T + m$  the cointegrating relation breaks down. The breakdown may be due to (i) a shift in the cointegrating vector from  $\beta_0$  to  $\beta_t$ , (ii) a shift in the distribution of  $u_t$  from being stationary to being a unit root process, (iii) some other shift in the distribution of  $\{u_{T+1}, \dots, u_{T+m}\}$  from that of  $\{u_1, \dots, u_m\}$ , or (iv) some combination of the previous shifts. Note that the set up does not require the break to occur exactly at  $T + 1$ , but rather in the interval  $\{T + 1, \dots, T + m\}$ .

To test for cointegration breakdown Andrews and Kim (2006) suggest the use of two tests which they denote by P and R. The authors have a slight preference for the P test because in a Monte Carlo simulation study this test showed somewhat better size properties than the R test. To determine the critical values and the p-values of the tests a parametric sub-sampling technique is used, as proposed by Andrews and Kim (2006), instead of large-sample asymptotics<sup>10</sup>.

For the models under scrutiny the tests are conducted under the assumption that cointegration and long-run stability hold when the models are estimated with data until the third quarter of 2001. Thus, cointegration breakdown is investigated for the period 2001Q4-2006Q4. The choice of this period stems from the fact that the second half of 2001 marks the beginning of a period of particularly high M3 growth. On the other hand the validity of the tests rests on the assumption that the model is stable before the date of the break and there is evidence that the models are stable when estimated with data until 2001Q3 (see Carstensen, 2004).

Table 2.2 presents the simulated p-values of the P and R tests for the three models under investigation, using FM-OLS and FIML to estimate the long-run relationships. From Table 2.2 we see that there are no strong signs of instability or cointegration breakdown in the three models when the sample until 2003Q2 is considered<sup>11</sup>. When the sample is extended until 2004Q4, cointegration and/or stability is generally rejected in the CGL models (the exception is the R test in the “revised version” estimated by FIML), but not in the Carstensen specification. However, when data until 2005Q4 and 2006Q4 are considered cointegration and/or stability of the three models is strongly rejected (the exception are the tests in the Carstensen specification estimated by FIML with the full sample period).

Thus, the evidence we get from the Andrews and Kim cointegration breakdown tests

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<sup>10</sup>The P and R tests in Andrews and Kim (2006) correspond to the Pc and Rc tests suggested in Andrews and Kim (2003) and reported in Alves et al. (2006). In the computations of the Andrews and Kim tests we used a Rats procedure developed by Carstensen for Rats 6.3, which we downloaded from ESTIMA webpage.

<sup>11</sup>This is the sample considered in Carstensen (2004, 2006).

**Table 3:** Cointegration breakdown tests (P-values)

Test	CGL (original)	CGL (revised)	Carstensen
Break at 2001Q4, sample until 2003Q2			
P (FM-OLS)	0.100	0.063	0.188
R (FM-OLS)	0.075	0.051	0.188
P (FIML)	0.115	0.182	0.205
R (FIML)	0.090	0.364	0.115
Break at 2001Q4, sample until 2004Q4			
P (FM-OLS)	0.000	0.000	0.135
R (FM-OLS)	0.000	0.000	0.243
P (FIML)	0.000	0.000	0.389
R (FIML)	0.000	0.211	0.528
Break at 2001Q4, sample until 2005Q4			
P (FM-OLS)	0.000	0.000	0.000
R (FM-OLS)	0.000	0.000	0.000
P (FIML)	0.000	0.000	0.029
R (FIML)	0.000	0.000	0.000
Break at 2001Q4, sample until 2006Q4			
P (FM-OLS)	0.000	0.000	0.000
R (FM-OLS)	0.000	0.000	0.000
P (FIML)	0.000	0.000	0.063
R (FIML)	0.000	0.000	0.234

Note: Entries in the Table are the bootstrapped P-values of the P and R cointegration breakdown tests proposed in Andrews and Kim (2006).

is in line with the evidence on the Johansen cointegration tests presented above. When data for the period 2004-2006 are considered, cointegration is progressively lost and stability is rejected both in the CGL and in the Carstensen models.

Acknowledging this fact, several recent studies attempted to find alternative specifications that purportedly yield long-run stable money demand functions for M3 in the euro area. These include Dreger and Wolters (2006), Landesberger (2007) and Greiber and Setzer (2007). However, these models also exhibit robustness problems and/or do not withstand the inclusion of the latest data<sup>12</sup>.

Overall, the results in this section show that a stable long-run relation linking M3 and its traditional determinants ceased to exist. This cointegration breakdown implies that the monetary indicators based on the residuals of the cointegrating regressions, such as the monetary overhang/shortfall and the real money gap, lose their interpretation as excess liquidity indicators and their information content with respect to future inflation<sup>13</sup>.

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<sup>12</sup>The model suggested in Dreger and Wolters (2006) involves the real stock of M3, real GDP and inflation. However, it seems that the model was not estimated using the official M3 aggregate but rather a monetary aggregate built on the basis of money holdings not adjusted for reclassifications, other re-evaluations, exchange rate variations and variations other than those related to transactions. When the official aggregate is used, the evidence clearly does not support the existence of a cointegrating relation.

The model suggested in Landesberger (2007) involves the real stock of M3, real GDP, the log of the yield on long-term government bonds and the dividend yield of the euro area equity markets and was identified for the period 1991Q1-2005Q4. However, when data for 2006 are added to the sample evidence on cointegration weakens significantly, the coefficient on the long-term interest rate changes its sign, the exogeneity restrictions imposed by the author are no longer accepted and the Andrews and Kim test clearly rejects the null of stability of the model, especially so if the model is estimated by Fiml.

The models suggested in Greiber and Setzer (2007) aim to take into account the role of the housing market in the behaviour of M3. At a general level, this is a very interesting contribution, and a step forward in the literature on money demand in the euro area. However, empirically, these models have serious shortcomings. The two models suggested by the authors involve the real stock of M3, real GDP, the yield on long-term government bonds and either the real residential property price or the housing wealth indicator. Both models were identified for the period 1981Q1-2006Q4 and the authors found strong evidence supporting a cointegrating relation for both models. However, we were unable to closely replicate the results with our dataset (which corresponds to the dataset officially used by the ECB). For the model that includes the real residential property price we get an estimate for the coefficient of GDP (0.089) which is much lower than the one presented by the authors (0.32) and moreover is not significantly different from zero. For the model that includes the housing wealth indicator we also get estimates for the long run parameters that are clearly different from those reported by the authors. In particular the coefficient of the ten year government bond yield is wrong signed and non-significantly different from zero. From the recursive estimates we conclude that statistical non significance of GDP, in the first model, and of the ten year government bond yield in the second, starts back in late nineties suggesting that the differences in the two data sets involve much more than simple end-of-sample revisions.

<sup>13</sup>A formal demonstration of the implications of cointegration breakdown for the interpretation and for the leading indicator properties of the excess liquidity indicators based on the residuals of the cointegrating regressions can be seen in Alves et al. (2006).

### 3 Stability of the Phillips-curve with filtered data

In this section we investigate the stability of Phillips-curve type models for the euro area that use filtered data of inflation and money growth<sup>14</sup>. This type of models has been used to test the idea that there may be a close relationship between inflation and money growth in the long-run, but that such a relation does not hold or does not need to hold in the short-run. Such relations, which have been termed “Two-Pillar Phillips curves” in Assemacher-Wesche and Gerlach (2006a, 2006b), may be estimated either by using pre-filtered series (which are seen as estimates of the long-run or low frequency components of the series) or by resorting to band spectral estimators suggested in Engle (1974) and Phillips (1991), which integrate the filtering and estimation techniques. Examples of these two estimation procedures can be seen in Neumann and Greiber (2004) and Assemacher-Wesche and Gerlach (2006a, 2006b).

In this section we investigate the stability of the Phillips-curve specification suggested in Assemacher-Wesche and Gerlach (2006a) which may be written as

$$\pi_t = \alpha_g g_{t-1} + \{ \alpha_m \Delta m_t^{LF} + \alpha_y \Delta y_t^{LF} + \alpha_\rho \rho_t^{LF} \} + \varepsilon_t \quad (4)$$

where  $\pi_t$  stands for current inflation,  $g_t$  for the output gap and  $\Delta m_t^{LF}$ ,  $\Delta y_t^{LF}$  and  $\rho_t^{LF}$  for the long-run (i.e., the low-frequency) component of money growth, output growth and long term interest rate changes, respectively. According to the quantity theory of money demand one should have  $\alpha_m = 1$  and  $\alpha_y = -1$  in (4).

Equation (4) is a formalization of the idea that long-term variations in inflation are closely associated with long-term movements in money growth, GDP growth and nominal interest rate changes, while short-term (i.e., high frequency) fluctuations in inflation are correlated with the output gap. In addition to the output gap, cost-push factors (changes in import prices, in oil prices or in the nominal effective exchange rate) may also be considered in equation (4) to better explain short-term fluctuations in inflation (see Assemacher-Wesche and Gerlach, 2006b).

We have seen in section 2 that M3 money demand equations in levels exhibit instability/cointegration breakdown when data until 2006Q4 are used with the models estimated in the time domain. Cointegration of time series is a long-run property so that in the frequency domain it refers to the zero-frequency relationship of the series. Accordingly, there is a frequency-domain equivalent of the time-domain cointegration

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<sup>14</sup>In this section we use the expression “filtered series” to designate the trend or low frequency component of the original series rather than the cyclical component.

analysis<sup>15</sup>. However, the evidence in Section 2 on the cointegration breakdown tests for money demand regressions in levels does not necessarily carry over to equation (4). This is mainly for two reasons. First, because equation (4) is defined in differenced variables so that it may be the case that cointegration is present in (4) (if it still includes I(1) variables) despite being absent of the corresponding equation in levels. Second, in the money demand equations of section 2 the real money stock is modelled as a single variable so that in fact long-run price homogeneity is being imposed at the outset while in equation (4) the coefficient of  $\Delta m_t^{LF}$  is estimated freely (and may turn out not to be equal to 1). Thus, formally, the evidence in section 2 does not dispense us with the need to investigate the issue of instability/cointegration breakdown of equation (4).

For the euro area inflation and M3 growth have been found to be I(1) and GDP growth and long term nominal interest rate changes to be I(0) (see, among others, Assemacher-Wesche and Gerlach 2006a, 2006b). Thus, before proceeding it may be instructive to look at the relationship between money growth and inflation in the time domain, for which cointegration tests are available.

If we estimate a VAR involving inflation and nominal M3 growth for the period 1980Q2-2006Q4 we find that the two series are cointegrated and this conclusion does not depend on whether we compute inflation using the HICP series or the GDP deflator. When the HICP is used the recursive estimates of the coefficient on money growth increase steadily from 1.13 in 2001Q3 to 1.26 in 2006Q4 (from 1.07 to 1.28 in the case of the GDP deflator) but, despite this, the recursive graphics do not suggest strong stability problems in this period<sup>16</sup>. In turn, the Andrews and Kim (2006) tests do not reject the null of stability in the model with the HICP for the period 2001Q4-2006Q4<sup>17</sup> while for the same period stability is rejected for the model with the GDP deflator<sup>18</sup>. This outcome for the model with HICP appears at first sight as unexpected given that M3 growth has been drifting apart from inflation for most of the last five years of the sample, as can be seen in Figure 3.

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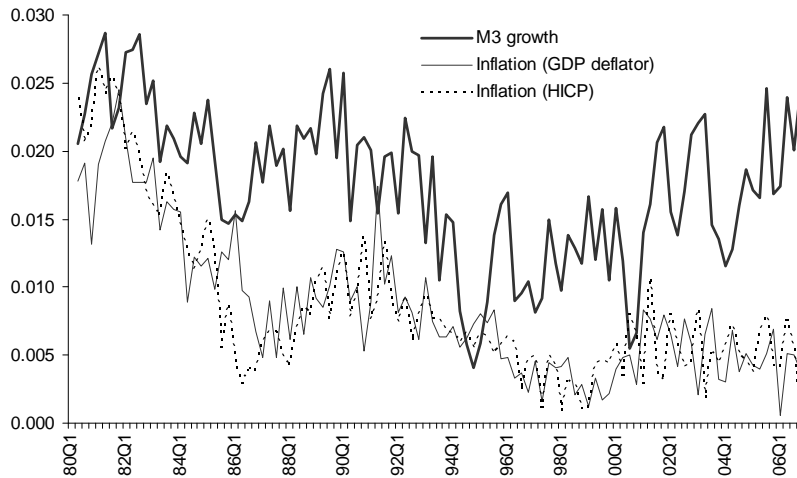
<sup>15</sup>Specifically, existence of a cointegration relationship between two time series in the time domain imposes restrictions on the zero-frequency behaviour of the series in terms of their cross-spectral measures in the frequency domain (coherence, phase and gain). For details, see Levy (2002). On the other hand, a regression model in the time domain can be transferred into the frequency domain by taking Fourier transforms of the data and estimating it on the transformed variables (see, Engle, 1974, for the stationary case and Phillips, 1991, for the nonstationary case). If all the frequencies are included, estimation of a cointegrating regression in the frequency domain is equivalent to the estimation in the time domain.

<sup>16</sup>We estimate the model with three lags in the case of HICP and two lags in the case of GDP deflator, with the constant restricted to the cointegration space.

<sup>17</sup>The p-values of the P and R tests are 0.23 and 0.33 if the model is estimated by FIML and 0.12 and 0.15 if estimated by FM-OLS.

<sup>18</sup>The rejection obtains at a 5 percent significance level for the P test and at 1 percent level for the R





**Figure 3:** Quarterly M3 growth and inflation

One likely explanation for the above findings stems from the fact that the high relative volatility of the quarterly M3 growth implies that the long-run coefficient on this variable is very imprecisely estimated. In particular, in the case of the model with HICP the standard deviation is 0.27 for a point estimate of 1.26 (when the full sample is used), which implies a very wide 95% confidence interval (0.73;1.79) for that coefficient. This, in turn, also explains why the restriction of a unit coefficient on money growth is comfortably accepted, in both models. As we will see below excluding the short-term (high frequency) fluctuations from the data has important consequences for the analysis.

We thus now turn to the Phillips-curve (4). In order to run the Andrews and Kim (2006) tests we need to estimate equation (4) using pre-filtered series so that the possibility of using the above mentioned band spectral estimator is excluded. In the exercise performed here we focus on the results when the long-run or low frequency component of the series is computed using the HP filter with  $\lambda = 1600$ , but as a robustness check we also look at alternative definitions of this low frequency component<sup>19</sup>.

Given the use of filtered series in (4) one may expect the residuals  $\varepsilon_t$  to be strongly autocorrelated so that in the estimation we use the FM-OLS procedure as it has the

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test, irrespectively of whether the model is estimated by FIML or by FM-OLS.

<sup>19</sup>It has been shown (see Kaiser and Maravall, 1999 and Pederson, 2001) that by using the HP-filter with  $\lambda = 1600$  one effectively filters out all the fluctuations with a frequency of less than 40, 36 or 32 quarters (10, 9 or 8 years) the exact value depending on the features of the data. This means that in this case we are defining the long-run as the frequency bands implying periodicities of 8-10 years to infinity.

advantage (over the FIML) of not requiring the specification of the precise model for the short-run dynamics and of being compatible with different types of error processes. Moreover, Li et al. (1995) using simulation methods investigated the properties of this estimator in cointegration equations involving HP trends of I(1) variables and concluded that the FM-OLS generally performs better than OLS. In turn, Li (1998) concluded that the FM-OLS on low-pass filtered data shares the same asymptotic efficiency as the FM-OLS on original data, but may gain efficiency in finite samples<sup>20</sup>.

We estimate equation (4) using both the HICP and the GDP deflator as alternative measures of inflation. In addition we also estimate (4) first unrestrictedly and then imposing the restriction suggested by the quantity theory that  $\alpha_m = -\alpha_y$ , such that M3 growth less output growth,  $\Delta m_t^{LF} - \Delta y_t^{LF}$ , enters as an explanatory variable. For presentation purposes we shall denote these two models the “unrestricted” and “restricted” versions of equation (4) and this new variable as “core M3 growth” (see Neumann and Greiber, 2004). The outcome of the Andrews and Kim (2006) tests for the four estimated models when the HP filter with  $\lambda = 1600$  is used to define the long-run are in Table 3<sup>21</sup>.

As in the previous section we consider the possibility of a break occurring after 2001Q3. As can be seen from Table 3, as more recent data are used in the analyses the evidence against the null of model stability generally increases. And, in particular, when data until 2006Q4 are used the null hypothesis of stability of the model is rejected in the four models<sup>22</sup>.

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<sup>20</sup>Note that this evidence does not contradict the results in Meyer and Winker (2005) which show that the “spurious regression” problem may be especially acute in regressions involving estimated HP trends of stationary variables. This “spurious regression” problem is not an issue in our case, however, because the tests in the time domain show that inflation and M3 growth are cointegrated and this, as remarked above, means that the regressions involving the low frequency components of these two variables are not spurious. On the other hand, our purpose here is to investigate the stability of the model at the end of the sample and not to perform significance tests on the estimated coefficients so that we are not using the t-statistics (nor the  $R^2$  statistic) to draw relevant conclusions. An open issue, however, is whether the use of filtered series in regression (4) may have a significant impact on the size and power of the Andrews and Kim tests.

<sup>21</sup>The models were estimated by FM-OLS using the Parzen window and assuming (i) a constant in the regression and (ii) that regressors do not show linear trend behaviour.

<sup>22</sup>The only exception regards the P test in the restricted model with the HICP (column 3). The fact that the p-values of the P test are usually higher than the p-values of the R test may be a reflection of some power problems of the P test. In fact according to the simulations in Andrews and Kim (2006) the best test in terms of power is the R, because it has a less variable power across different distributions than the P test.

Notice also that the p-values of the two tests are higher in the restricted model with the HICP (column 3) suggesting that this model may be closer to being stable. However this can hardly be seen as a positive outcome. In fact, when the full sample is used the estimates in the unrestricted model for the coefficients  $\alpha_m$  and  $\alpha_y$  are -15.299 and -21.243, respectively, and thus are far from complying with the imposed theoretical restriction  $\alpha_m = -\alpha_y$ . Moreover not only  $\alpha_m$  but also  $\alpha_\rho$  and  $\alpha_g$  are wrong signed. This casts strong doubts on the interpretation and usefulness of the model.

**Table 4:** Cointegration breakdown tests (P-values) for Two-Pillar Phillips curve

Test	HICP		GDP Deflator	
	Unrestricted	Restricted	Unrestricted	Restricted
Break at 2001Q4, sample until 2003Q2				
P(FM-OLS)	0.013*	0.150	0.013*	0.038*
R(FM-OLS)	0.013*	0.125	0.013*	0.013*
Break at 2001Q4, sample until 2004Q4				
P(FM-OLS)	0.135	0.270	0.014*	0.014*
R(FM-OLS)	0.014*	0.135	0.000**	0.014*
Break at 2001Q4, sample until 2005Q4				
P(FM-OLS)	0.286	0.343	0.229	0.029*
R(FM-OLS)	0.300	0.114	0.014*	0.014*
Break at 2001Q4, sample until 2006Q4				
P(FM-OLS)	0.000**	0.349	0.000**	0.015*
R(FM-OLS)	0.000**	0.015*	0.000**	0.015*

Note: Entries are the bootstrapped P-values of the P and R tests proposed in Andrews and Kim (2006). \* and \*\* mark rejection of stability or cointegration at 5% and 1%, respectively.

The outcome of the Andrews and Kim (2006) tests may however be dependent on the way the trend of the variables is defined. Thus, as a robustness check we conducted stability tests using alternative measures for the trend or low frequency components of the variables in equation (4). In particular, using the nonsymmetric version of the Christiano-Fitzgerald filter we computed four alternative definitions for the long-run or low frequency components of the series. These are defined such that they include all the frequencies implying periodicities between 16, 24, 32 or 40 quarters (4, 6, 8 or 10 years) and infinity, respectively<sup>23</sup>.

The results for the different models corroborate the conclusions obtained with the HP filter in Table 3. When the long-run is defined such that it includes frequencies implying periodicities larger than 24, 32 or 40 quarters (in the unrestricted versions of the models for both the HICP and the GDP deflator) the null of stability is always rejected at a 5% significance level<sup>24</sup>. Only when the long-run is defined such that it includes all the

<sup>23</sup>The Christiano-Fitzgerald filter was computed assuming that inflation is I(1) and that GDP growth and nominal interest rate changes are both I(0).

<sup>24</sup>For two cases (long-run including periodicities larger than 24 or 32 quarters) it was not possible to test stability for the usual period 2001Q4-2006Q4, because due to strong multicollinearity problems the FM\_OLS method failed to compute the relevant statistics. However for those cases stability for the period 2005Q1-2006Q4 is clearly rejected.

frequencies implying periodicities between 16 quarters and infinity do the Andrews and Kim (2006) tests fail to reject the null of stability for a 5% test<sup>25</sup>.

In summary the Phillips-curve involving inflation, trend M3 growth, trend GDP growth, trend nominal interest rate changes and the output gap has strongly deteriorated in recent years such that the null hypothesis of stability of the model is generally rejected when data until 2006Q4 are used in the estimation. This implies that the empirical support for the two-pillar Phillips curves proposed in Assemacher-Wesche and Gerlach (2006a, 2006b) fades away when the most recent data are taken into account.

## 4 M3 as a leading indicator of prices in the euro area

The leading indicator properties of money regarding future prices have always been regarded as a centrepiece of the special role assigned to monetary analysis in the ECB's monetary policy strategy. This feature is grounded on the property of long-run monetary neutrality and on the fact that the transmission mechanism from changes in money to subsequent changes in prices occurs with a significant lag (Friedman, 1970 and Lucas, 1980). This latter feature has been previously investigated for the euro area, and empirical studies concluded overall that M3 growth was a good predictor of future inflation (Altimari, 2001 and Trecroci and Vega, 2002). However, these studies preceded the strong dynamics of M3 in recent years.

Against this background, this section aims at documenting the leading indicator properties of M3 including the most recent evidence. Given the cointegration breakdown in the long-run M3 money demand function, the leading indicators properties of the M3 aggregate and of the excess liquidity indicators computed using this aggregate would be expected to deteriorate significantly. This would imply that M3 would no longer be a good instrument to analyse the medium to long-term prospects for inflation in the euro

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<sup>25</sup>This result could be expected given the evidence above where stability in the time domain was obtained for a model involving money growth and inflation (using HICP). As higher frequency components (shorter term trends) are included in the definition of the long-run component of the series the evidence against the null of stability may be expected to weaken because local flexibility is increasing and thus the difference between the defined trends and the actual series is diminishing. Notice however that in this version of the model the coefficient of GDP is wrong signed, so that stability of the model can hardly be seen as a positive outcome.

In order to check whether instability (and wrong signed parameters) could be due to strong multicollinearity problems we also carried out the stability tests by estimating model (4) without the interest rate component,  $\rho_t^{LF}$ , as this variable exhibits very low variability such that in most cases the associated parameter is not significantly different from zero. With the exclusion of  $\rho_t^{LF}$  most previously wrong signed coefficients turn out to display the right sign, but for all the estimated models the null of stability for the period 2001Q4-2006Q4 is clearly rejected, including the ones for which the long-run is defined such that it includes all the frequencies implying periodicities between 16 quarters and infinity.

area.

To analyse this issue using the most recent data, we undertake two different types of exercises. In subsection 4.1, we evaluate the performance of M3 in a simulated out-of-sample forecast exercise, following Altimari (2001). In subsection 4.2, we analyse frequency-wise measures of causality, following the procedure proposed in Breitung and Candelon (2006).

#### 4.1 Results based on the Stock and Watson (1999) approach

In order to assess the existence of a leading indicator role for M3 growth with respect to inflation in the euro area Altimari (2001) applied the methodology proposed by Stock and Watson (1999) to the euro area. This methodology compares the forecast performance of univariate models of inflation with that of bivariate models including M3 growth as an additional explanatory variable. According to the results of Altimari (2001), M3 growth has leading indicator properties for inflation in the two to three year-ahead horizons. This conclusion is based on specifications which assume that M3 growth and inflation are stationary variables during the sample period<sup>26</sup>.

Assuming both M3 growth and inflation are I(0) the approach involves assessing the forecast performance of the following model of inflation:

$$\pi_{t+h}^h = c + \gamma(L)\pi_t + \omega(L)\Delta M3_t + u_{t+h} \quad (5)$$

where the annualised inflation rate over the following  $h$  quarters  $\pi_{t+h}^h = (4/h) \ln(P_{t+h}/P_t)$  is modelled as a function of current and lagged values of annualised quarterly inflation  $\pi_t = 4 \cdot \ln(P_t/P_{t-1})$  and lagged values of the growth rate of M3  $\Delta M3_t = 4 \cdot [\ln(M3_t) - \ln(M3_{t-1})]$ .  $\gamma(L)$  and  $\omega(L)$  are polynomials in the lag operator  $L$  and  $u_{t+h}$  is the out-of sample forecast error. The number of lags (which can range from 0 to 3) is selected using the Schwartz criterion. The exercise starts by estimating model (5) from 1980Q1 until 1996Q4 and computing out-of-sample forecasts for horizons  $h$  varying from one quarter to three years ahead. The sample is then recursively extended quarter by quarter, and each time out-of-sample forecasts are computed for the various horizons  $h$ . For each horizon, the (average) ratio of the root mean square forecast error of models with M3 growth over that of a univariate model of inflation can then be computed, for

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<sup>26</sup>This hypothesis is, however, rejected by the data, which suggest that both inflation and money growth are better classified as integrated of order 1. When the specification takes into account the properties of the series in the sample period, the information content of monetary aggregates tends to disappear even in the period 1997-2003. This finding is reported in Altimari (2001), but usually is not duly emphasised in the quotations of the paper.

**Table 5:** Forecast performance of M3 in bivariate models of inflation

Horizon	1	2	3	4	5	6	7	8	9	10	11	12
HICP												
1997Q1-2006Q4	1.1	1.1	1.1	1.1	1.0	0.9	1.0	1.0	1.3	1.3	1.0	0.7
1997Q1-2001Q3	1.4	0.9	0.8	0.9	0.5	0.5	0.6	0.6	1.0	0.9	0.6	0.4
2001Q4-2006Q4	0.9	1.2	1.4	1.2	1.8	1.8	2.0	2.2	1.8	2.3	2.5	2.9
GDP deflator												
1997Q1-2006Q4	1.2	1.2	1.6	1.6	1.5	1.3	1.1	1.1	1.0	1.0	0.9	0.7
1997Q1-2001Q3	1.6	2.5	1.9	1.1	0.8	0.6	0.5	0.5	0.5	0.4	0.3	0.3
2001Q4-2006Q4	1.1	1.1	1.5	2.2	2.5	2.5	2.1	2.0	2.1	2.2	2.3	2.5

Note: models assume M3 growth and inflation are stationary variables.

any sub-sample between 1997Q1 and 2006Q4.

Table 4.1 shows the results of this exercise, using the HICP and the GDP deflator. The results are presented for the period 1997Q1-2006Q4, and for two subsamples, before and after 2001Q3. There are several conclusions that deserve being highlighted. First, the table confirms the results in Altimari (2001) that in the period before 2001Q3 the bivariate model using M3 growth performs better than the univariate model, particularly at longer horizons. Second, the table also shows that in the most recent sample period, the leading indicator properties of M3 growth vanish entirely. This confirms that the breakdown of M3 money demand was concurrent with a significant deterioration of the leading indicator properties of M3. Third, the results are robust to the choice of inflation indicator. Finally, we confirmed (in results not shown in the table) that M3 does not provide additional information to forecast inflation in the full sample when both inflation and money growth are assumed to be non-stationary, a result that mirrors the original Altimari (2001) study.

Overall it can be concluded that the favourable forecasting performance of M3 observed up to 2001 has noticeably deteriorated since then. In this context, it should be noted that the most recent monetary dynamics are by construction excluded from the evaluation of the forecasting performance of the bivariate monetary models. For example, when assessing the performance of the models in the 12-quarter ahead horizon, the M3 data ends in 2003Q4. Given the accelerating pattern of M3 in the most recent quarters and the current projections for inflation at levels broadly in line with the ECB's objective of price stability, it is straightforward to conjecture that the performance of M3 will deteriorate even further in the near future when evaluated through

this methodology.

## 4.2 Causality tests in the frequency domain

A formal way of testing the leading indicator properties of money in the frequency domain is to resort to statistical tests, such as the one proposed by Breitung and Candelon (2006). The idea is that M3 may not be informative about the high frequency components of inflation but may have predictive content for trend inflation. In order to conduct this test, a bivariate system containing quarterly M3 growth and inflation (computed using the HICP) is set-up using the sample 1980Q1-2006Q4. The test is based on a regression of inflation on lagged values of inflation and money growth:

$$\pi_t = c + \gamma(L)\pi_t + \beta(L)\Delta M3_t + u_t \quad (6)$$

The causality test then consists of assessing whether the coefficients on lagged money growth are statistically significant in an equation of inflation on past inflation and money. The hypothesis that money Granger causes inflation at frequency  $\omega$  is equivalent to testing whether the following linear restriction holds:

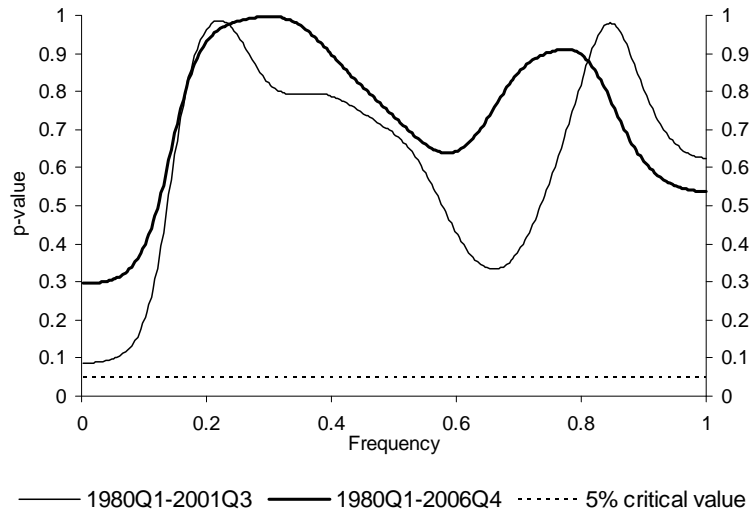
$$H_0 : R(\omega)\beta = 0 \quad (7)$$

where  $R(\omega) = [\cos(\omega) \quad \cos(2\omega) \quad \dots]$ . The test statistic is approximately  $F(2, T - 2p)$  for  $\omega \in (0, \pi)$ .

In order to select the appropriate lag length of the VAR we rely on the Akaike information criterion, which is minimised for a lag length of 8 quarters.

The test is conducted for different frequencies and the results are shown in the Figures below. As can be seen in Figure 4, there is no evidence that M3 growth causes inflation for any frequency when the sample used goes from 1980Q1 to 2006Q4 (the horizontal line show the 5% critical values for the null hypothesis that the coefficients on lagged money growth are zero). However, if one stops at 2001Q3, then there is evidence that M3 growth has predictive content for the long-run trend in inflation, but only at a 10% significance level.

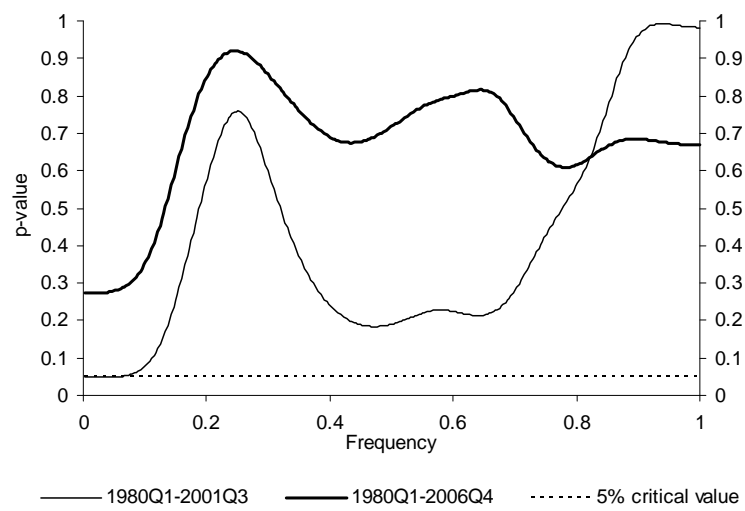
These conclusions also hold if, as done by Assenmacher-Wesche and Gerlach (2006a), one includes the output gap and changes in the long-term rate as conditioning variables (Figure 5). In this case, the evidence supporting causality from M3 growth to the low frequency movements in inflation in the period until 2001Q3 is even clearer. However, for the full sample period, one concludes that the causality from M3 growth to inflation



**Figure 4:** Frequency domain causality test: M3 growth causing inflation.

ceases to be significant, for any frequency. Thus, this analysis seems to be in line with the results of the previous section which suggest a deterioration of the information content of M3 in recent years.





**Figure 5:** Frequency domain causality test: M3 growth causing inflation, including also the output gap and changes in the long-term rate.

## 5 Conclusions

This paper reassesses several empirical properties of the M3 monetary aggregate in the euro area and discusses several implications for monetary policy purposes. The analysis carried out in the paper allows us to conclude that the available M3 money demand models show strong signs of instability or cointegration breakdown when data up to the end of 2006 are considered. The emergence of cointegration breakdown has the implication that a stable long-run relation linking real M3, the level of activity and opportunity costs, as specified in those models, ceased to exist. Cointegration breakdown also implies that the monetary indicators based on the residuals of cointegrating regressions of such models, as the monetary overhang/shortfall and the real money gap, lose their interpretation as excess liquidity indicators. The analysis in the paper also shows that the stability of the two-pillar Phillips curves involving inflation and trend money growth is generally rejected when the most recent data are added to the sample, and that the leading indicator properties of M3 regarding inflation have severely deteriorated in recent years.

The evidence in this paper does not preclude the possibility of a new stable money demand relation to be found in the euro area, namely by properly redefining the relevant aggregate or by duly accounting for the implications of deregulation and financial innovation, as well as the existence of cross-border portfolio flows in an open economy framework. On the other hand, the possibility of cointegration breakdown in M3 money demand equations being related to structural reasons cannot be dismissed. Most importantly, the introduction of the euro may have represented a true regime shift in the sense of Lucas (1976), namely in what concerns the degree of monetary and financial integration. This structural change, coupled with the ongoing innovation and deepening in global financial markets, should be expected to have a significant impact on the behaviour of economic agents in the euro area, in particular on their portfolio decisions (Papademos, 2007).

It may also be noted that the demise in the information content of M3 is not surprising, from an historical perspective. Over the last decades, and across numerous countries, monetary aggregates with solid signalling properties with respect to output and/or prices have started displaying such low levels of signal to noise that led monetary authorities to abandon them (see Calza and Sousa, 2003, and Friedman and Kuttner, 1996). In the famous quote from former Bank of Canada governor Gerald Bouey “We didn’t abandon the monetary aggregates; they abandoned us”. These developments have been usually associated with processes of deregulation and financial innovation, in

particular concerning the technology underlying the transactions role of money.

The evidence in this paper challenges the two properties of the M3 aggregate that justified its prominent role in the ECB's monetary analysis. However, it must be underlined that such evidence does not imply that monetary analysis - interpreted in a broad sense, including the behaviour of various monetary aggregates, counterparts, country/sectoral breakdowns and households and firms' balance sheets - is not useful. On the one hand, the monetary analysis is indispensable in order to capture important channels in the transmission mechanism of monetary policy, in particular in the presence of financial frictions (see Galí et al., 2004, King, 2002, Bernanke et al. 1999, and Goodfriend and McCallum, 2006). On the other hand, monetary variables may provide information that is helpful for understanding the state of the economy, aggregate demand, asset prices and financial conditions (see Nelson, 2002, Dotsey and Hornstein, 2003, Dotsey et al., 2000, Coenen et al., 2005 and Machado and Sousa, 2006).

## A Variables used in the money demand models (sub-sample 1990Q1-2006Q4)

Figure A1

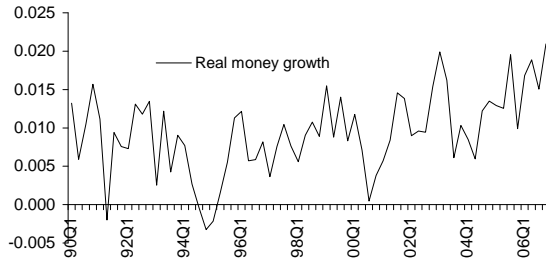


Figure A2

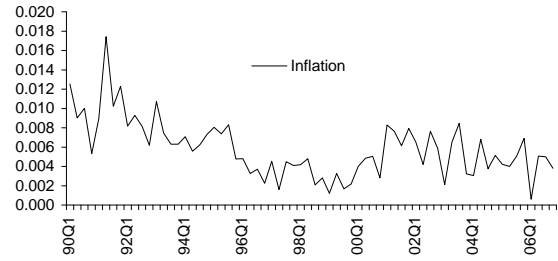


Figure A3

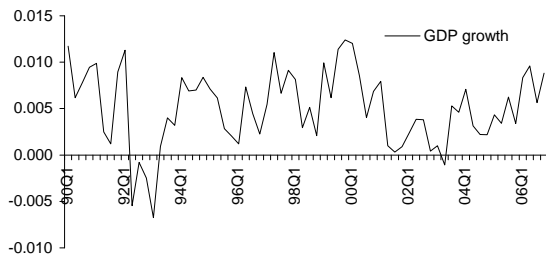


Figure A4

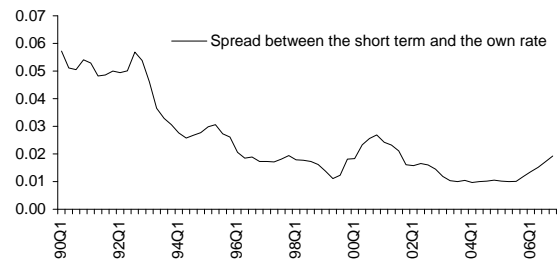


Figure A5

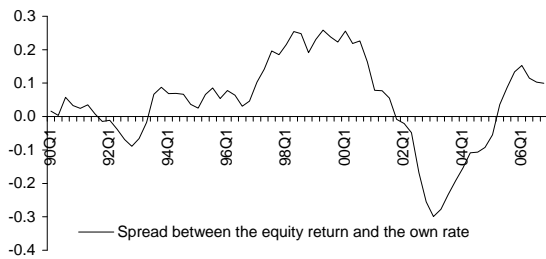
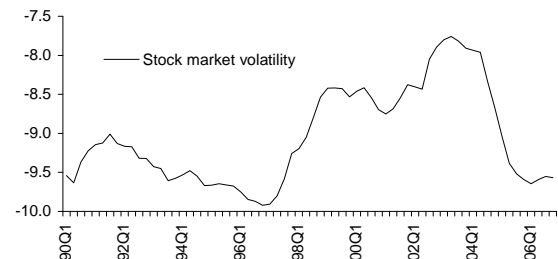


Figure A6



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