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Which of the real money gap or nominal money gap helped predict inflation in Europe? A retrospective analysis

Abstract

The question examined in this paper is the following. Assuming that money played a role in the prediction of inflation, which of the *nominal* money gap or *real* money gap did the best job in the European countries? Answering this question helps us to compare the different strategies undertaken by the central banks in the countries that were members of the EMU. In the countries that participated in the Exchange rate mechanism (ERM) and then adopted the Euro, the policy preferences have been dominated by tacking monetary aggregates, while some non-euro countries preferred to focus on the direct effects of real money growth. The authors use panel data econometrics allowing for heterogeneous short-run and long-run dynamics among the countries. An important result is that the real money gap may be equally informative about future inflation. This plays against the dominant view of a quantitative theory approach of inflation in Europe.

Keywords: monetary policy, inflation, panel data.
JEL Classification: E42, E52, E58, C23, C53.

Introduction

Since the birth of the Economic and Monetary Union (EMU), medium-to-long-term orientation of monetary policies has implied the achievement of price stability. Whether or not money has been a reliable indicator for inflation in the European countries has been a matter of debate. On the one side, some authors criticized the money pillar approach of monetary policy pointing to the lack of theoretical foundations and arguing that all relevant information for money is already included in past inflation and that there is no role for monetary aggregates in any form (see for instance Woodford, 2008). On the other side, there is strong evidence in the empirical literature of a link between money growth and inflation in the euro area (see, among others, Benati, 2009; Hofmann, 2008; Kaufmann and Kugler, 2008).

The objective of our paper is not to explore once again the predictability ability of money for inflation in the European countries. Our attention is rather on another aspect of the money-inflation link, which has been discarded although it has important implications in terms of monetary strategy. Our question is the following. Assuming that money played a role in the prediction of inflation, which of the *nominal* money gap or *real* money gap did the best job in the European countries? The point we want to make is that there has always been an implicit consensus in the European central bankers' mind that inflation pressures were mainly a monetary phenomenon. Indications coming from the academic literature by leading macroeconomists reinforced this view[1]. However, the view according to which inflation also resulted from the balance of supply and demand in other markets (goods and services, factors, financial) and that desequilibria in these markets were translated into money gaps should not be disregarded[2]. Despite the fact that the EMU is made of countries sharing and not sharing a common currency, there seems to be "a bias" towards the quantity theory approach of inflation in Europe. In spite of the fact that the European Central Bank (ECB) has assigned a prominent role to tracking the growth rate of nominal money, which has been designed as the second pillar of its monetary policy strategy, real money gap might be equally important for predicting inflation in non-euro area countries that are members of the EMU. The empirical analysis of this paper compares the forecasting performance of both indicators (nominal and real money gaps) over the past years from 1990 to 2004 (reasons for ending in 2004 is the instability in euro area after that time). It is shown that for non-euro area countries, the real money gap serves as a good indicator to help predict inflation, with a forecasting performance similar to models using nominal money gaps.

This result can help us understand the different options undertaken by the central banks in the countries that were members of the EMU. In the countries that participated in the Exchange rate mechanism (ERM) and then adopted the Euro, the policy preferences have been dominated by the Bundesbank's and aimed at establishing anti-inflation reputation. To achieve inflation credibility and anchor inflation expectations, nominal monetary targeting (by reacting to deviations of M3 growth from the reference value of 4.5%) appeared as the best way to satisfy both the objectives of simplicity and transparency. But during the same years, other countries have also succeeded in controlling long-

We thank Michael McLennan for excellent research assistantship, particularly for his help in inputting the data.

[1] See, for instance, Estrella and Mishkin (1997), Svensson (1999), Stock and Watson (1999).

[2] Empirical evidence in favor of the role of the real money gap as a determinant of inflationary pressures in the euro area is also found (see Nelson, 2000; Trecoci and Vega, 2000; Gerlach and Svensson, 2003; Neuman and Greiber, 2004).

run inflation by following an alternative policy that was a form of inflation targeting. The idea was that the central banks needed to have a complete model of the determinants of inflation because monetary policy could not rely on the tacking of intermediate monetary aggregate since the demand for money exhibited an instability in the context of financial innovation. For instance, during its mandate, the former Bank of England governor, Eddie Georges, mentioned nominal money only one time out of 29 speeches over two years[1].

We begin in 1990, a few years before the adoption of the Maastricht Treaty. We do not go too far in the past in order to avoid periods where monetary policies in Europe have been characterized by fiscal dominance. We end in 2004, because after that year inflation expectations have been characterized by greater uncertainty and this resulted in a higher persistence in the dynamics of inflation[2]. Such persistence could mask the potential influence of the macroeconomic variables that determine the money gaps.

We use panel data instead of the usual time series methods on aggregate Euro data. We allow for heterogeneous short-run and long-run dynamics among the European countries. To bring evidence regarding the respective predictive power of the real money gap and a nominal money growth indicator for future inflation, we follow two approaches. We first consider an inflation equation containing the real money gap as a monetary indicator, along with two other macroeconomic variables that potentially influence future inflation: the output-gap and a component reflecting the inflation expectations. To model the private sector inflation expectations, we assume that the agents align their inflation forecasts on the central bank's implicit inflation objective. Then, this model is compared to another one where, instead of the real money gap, we consider the growth of the nominal money as a monetary indicator.

The rest of the paper is organized as follows. Section 1 briefly delineates the theoretical models of inflation and real money demand. Section 2 presents our estimates of the inflation models. Section 3 compares the predictive power of the real money gap and a nominal money growth indicator using out-of-sample forecasts. The final sections presents the conclusion and the implications.

1. Inflation and real money demand equations

1.1. Inflation model. Inflation dynamics is described by a standard aggregate supply equation in which the first-difference of the general price level is a function of past inflation, demand pull inflation and inflation expectations:

$$\pi_{it} = \mu_1 \pi_{it\text{-}1} + \mu_2 \left(y_{it} - y_{it}^{*} \right) + \mu_4 \pi_{i,t,t\text{-}1}^{e} + \omega_{it}. \quad (1)$$

The index i refers to a specific country, while the index t refers to a quarter. All the variables written in lowercase are in logarithm. $\pi_{it} = 4\ (p_{it} - p_{it\text{-}1})$, where p_{it} is the general price level and π_{it} is the annualized inflation rate in quarter t, y_{it} is the real GDP, $\pi^{e}_{i,t,\ t\text{-}1}$ is the expectation in quarter $t-1$ of inflation in quarter t by the private sector in country i, $(y_{it} - y^{*}_{it})$, is the output-gap. ω_{it} is a disturbance term which can be interpreted as a supply-shock.

We need to make some hypotheses about the way inflation expectations are formed. We assume a backward-looking mechanism and also consider that they are influenced by the implicit inflation objective set by the monetary authority:

$$\pi^{e}_{i,t,t\text{-}1} = \overline{\pi}_{it\text{-}1} + c(\pi_{it\text{-}1} - \overline{\pi}_{it\text{-}1}),\ 0 \le c \le 1, \quad (2)$$

where $\overline{\pi}_{it}$ is the implicit monetary authorities' inflation target. As a proxy of this target, we consider the "stable" component of past inflation computed by an HP filter. In the sequel, we call this variable the core inflation, and the difference $\pi_{it} - \overline{\pi}_{it}$ the inflation gap in quarter t for country i.

We also consider that the output-gap is endogenously determined as follows:

$$y_{it} - y^{*}_{it} = \upsilon_1 \left(y_{it\text{-}1} - y^{*}_{it\text{-}1} \right) + \upsilon_2 \left(R^{c}_{it\text{-}1} - \pi_{it\text{-}1} \right) + \upsilon_3 \left(\hat{\varepsilon}_{it\text{-}1} \right) + u_{it}, \quad (3)$$

where R^{C} is the Central Bank's interest rate which is set through a Taylor rule:

$$R^{c}_{it} = \alpha_1 \left(\pi_{it} - \overline{\pi}_{it} \right) + \upsilon_{it}. \quad (4)$$

We assume that the Central bank only accommodates changes in inflation (not in the output gap) as has been the case of the European Central Bank.

Equation (3) is an aggregate demand equation in which output and the output-gap are used interchangeably as long as equilibrium output changes are not subject to structural breaks or brutal variations but vary smoothly (which was the case in the European countries over the period under examination). We introduce the role of monetary policy and consider several channels through which it can affect the activity. The first is the standard (real) interest rate channel which affects both consumption and investment spending. We further extend the conventional aggregate demand equation to incorporate a liquidity shock. $\left(\hat{\varepsilon}_{it-1} \right)$ is the real money gap defined as the

[1] The reader can also refer to the stimulating paper by Nelson (2000) from the Bank of England arguing for direct effects of real money growth.
[2] See Fountas et al. (2004), Conrad and Karanasos (2006).

difference between the real money stock and the long-run equilibrium real money stock. This effect can be introduced in models in order to capture cash-in-advance constraints in the good markets, the indirect influence of monetary policy when private demand is insensible to the interest rate variation, or wealth effects[1]. To account for a liquidity effect on aggregate demand, we need a definition of money as broad as possible (the indicator available for all the countries is M3).

After re-parameterization, the combination of (1) through (4) yields the following inflation equation:

$$\pi_{it} = \beta_1 \pi_{it\text{-}1} + \beta_2 \left(y_{it\text{-}1} - y^*_{it\text{-}1} \right) + \beta_3 \left(\pi_{it\text{-}1} - \overline{\pi}_{it\text{-}1} \right) + \\ + \beta_4 \left(\hat{\varepsilon}_{it\text{-}1} \right) + \beta_5 \overline{\pi}_{it\text{-}1} + \eta_{it}. \tag{5}$$

Where η_{it} is a disturbance term.

We further consider an alternative specification of equation (5) in which we substitute a nominal money growth gap indicator for the real money gap. In Equation (3), we replace $\hat{\varepsilon}_{it-1}$ by $\Delta m_{it\text{-}1} - m^*_{it\text{-}1}$, where Δm_{it} is defined as the annualized nominal money growth (similarly to the definition of the annualized inflation rate): $\Delta m_{it\text{-}1} = 4\ (m_{it\text{-}} - m_{it\text{-}1})$. This accounts for the so-called "direct money channel" according to which nominal money directly affects the output. The theoretical argument is that open market operations by a central bank, not only affects its nominal interest rate, but also its balance-sheet and thus the monetary base (which in turn implies changes in the growth rate of money supply, for instance through the credit multiplier). The target value of the nominal money growth, Δm^*_{it} is computed as the fitted values (static forecasts) of the following equation:

$$\Delta m_{it} = \delta_1 \pi_{it\text{-}1} + \delta_2 \left(y_{it\text{-}1} - y^*_{it\text{-}1} \right) + u_{it}. \tag{6}$$

This equation specifies a target for the growth rate of nominal money. Given inflation, a higher output-gap results from an expansionary monetary policy ($\delta_2 > 0$). But if money grows too rapidly above its target, then it may ultimately lead to high levels of inflation. The rate of nominal money growth thus would need to be adjusted downward ($\delta_1 > 0$) our equation.

To find out whether the real money-gap or a nominal money growth based indicator is best to predict inflation we shall compare (in section 4) the forecasting performances of equation (5) using those indicators as explanatory variables.

[1] Our model is not in contradiction with an ancient literature on "real balance effect" of the kind described by authors like Scitovsky, Haberler, Pigou or Patinkin. For a survey on the different arguments for introducing the real balance effect in macro models, we refer the reader to Ireland (2001).

1.2. A model of real money demand. To make our inflation model operational we need a model to determine long-run real money demand. Our long-run equation represents a standard model of the demand for real money:

$$m_{it} - p_{it} = \alpha^i_0 + a_1 \pi_{it} + a_2 y_{it} + a_3 R^L_{it} + a_4 R^s_{it} + \\ + a_5 RER_{it} + a_6 R^F_{it} + \varepsilon_{it}, \tag{7}$$

where m_{it} is the nominal money stock, p_{it} is the price level, $p_{it\text{-}} = 4\ (p_{it\text{-}} - p_{it\text{-}1})$, is the annualized inflation rate in quarter t, $y_{it\text{-}}$ is the real GDP, R^L_{it} is the long-run interest rate, R^S_{it} is the short-run interest rate, RER_{it} is the real effective exchange rate, R^F_{it} is the foreign interest rate and ε_{it} is a disturbance term. All the variables written in lowercase are in logarithm, while the remaining others are in level; ε_{it} is independently and identically distributed among countries and quarters.

Theoretically, we expect the coefficients to have the following signs. The influence of inflation must be signed negatively ($a_1 < 0$), since an increase in inflation means a higher return on holdings of real assets. This is likely to induce a substitution between money and goods. The interest rate variables capture the impact of financial asset substitution. We expect the long-run interest rate to carry a negative sign ($a_3 < 0$) and the short-run rate to be positive ($a_4 < 0$). A rise in the long-run interest rate will lead to a decrease of the demand for money (to take advantage of higher returns on bonds), while an increase in the short term interest rate will result in a higher demand for money. A rising foreign interest rate is likely to translate into a decrease in the money demand ($a_6 < 0$), caused by a propensity to substitute away from domestic money. We expect the coefficient of real GDP to be positively signed ($a_2 < 0$). Finally, an appreciation of the real effective exchange rate results in an increase in the demand for domestic currency, so that we expect the coefficient a_5 to be positive (because an appreciation of the domestic currency is reflected by an increase in *RER* – see the description of the data below).

2. The empirical models

2.1. The data. We use quarterly panel data spanning from 1990-1991 to 2004-2011. The empirical analysis focuses on the following group of the European countries: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Italy, the Netherlands, Norway, Portugal, Spain, Sweden and the United Kingdom. The data are from several sources: the International Monetary Fund IFS database, the national central banks' statistics, the OECD main indicators and Eurostat database. The series are seasonally adjusted.

The nominal money stock, m, is computed as the logarithm of M3. The price level, p, is measured as the logarithm of the consumer price index. The inflation rate, π, is computed as the annualized inflation (that is $\pi_{it} = 4\ (p_{it-} - p_{it\text{-}1})$. The long-run interest rate, R^L, is chosen as one of the following two variables (depending upon data availability): the Government bond yield or the bill rate. For the short-run interest rate, we use the money market rate as a proxy, or the commercial banks deposit rate when the money market rate series is not available. The real GDP, y, is the logarithm of the GDP at 1995 prices[1]. As a proxy for the foreign interest rate, we choose the US bond yield. Finally, the real effective exchange rate is measured as the trade weighted average of the real exchange rate based on bilateral trade shares. An increase in *RER* reflects an appreciation of the domestic currency. All the models are estimated over the period from 1990:1 to 1998:4. We use the data spanning from 1999:1 to 2004:1 to compute out-of-sample forecasts.

2.2. Empirical results for the real money demand model. To avoid the problem of spurious regressions, we begin our empirical investigation with an analysis of the unit root properties of the individual series. To this effect, we apply two tests based on Choi (2004) and Phillips and Sul (2003). They have some advantages over the conventional Im, Pesaran and Shin (2003) test (IPS) since they allow for cross-section dependence. The results in Table 1 indicate that we often conclude in favor of the unit root hypothesis, except for the inflation rate that is *I* (0).

Table 1. Panel unit root tests

		Choi(1)				Phillips and Sul (*Z*- test)(2)		
		Level	1st diff	concl		Level	1st diff	concl
	P_m	-0.293	8.56	I (1)	c	-1.73	-3.35	I (0)
Real money	Z	0.728	-5.74	I (1)	c,t	-0.35	-2.37	I (1)
	L^*	1.238	-6.26	I (1)				
	P_m	-0.88	3.85	I (1)	c	-3.33	-2.23	I (0)
Real GDP	Z	2.05	-3.48	I (1)	c,t	0.06	-0.38	I (2)
	L^*	2.32	-3.29	I (1)				
	P_m	4.67	-	I (0)	c	$-\infty$ (3)	-	I (0)
Inflation	Z	-2.96	-	I (0)	c,t	$-\infty$ (3)	-	I (0)
	L^*	-3.33	-	I (0)				
	P_m	-0.277	16.68	I (1)	c	-0.55	-5.03	I (1)
Long-term interest rate	Z	0.833	-9.49	I (1)	c,t	2.08	-3.30	I (1)
	L^*	0.94	-11.21	I (1)				
	P_m	-0.633	8.29	I (1)	c	0.99	-2.59	I (1)
Short-term interest rate	Z	0.66	-6.24	I (1)	c,t	1.29	-1.45	I (2)
	L^*	0.68	-6.41	I (1)				
	P_m	-0.27	16.68	I (1)	c	0.253	-6.97	I (1)
US bond yield	Z	0.833	-9.49	I (1)	c,t	2.01	-5.81	I (1)
	L^*	0.943	-11.21	I (1)				
	P_m	-0.10	4.88	I (1)	c	-1.42	-4.11	I (1)
RER	Z	0.109	-4.59	I (1)	c,t	-2.00	-1.74	I (0)
	L^*	0.05	-4.44	I (1)				

Notes: (1) All the statistics are distributed as standard normal asymptotically. The null hypothesis of a unit root is rejected for large positive values of the P_m statistic, while it is rejected for large negative values of the other two statistics. Accordingly, at the 5% level, we conclude as follows:

$$No\ unit\ root\ if \begin{cases} P_m > +1.64 \\ Z < -1.64 \\ L^* < -1.64 \end{cases}$$

(2) The statistic is distributed asymptotically as standard normal. The null hypothesis of a unit root is rejected for large negative values of the *Z*-statistic. We thus conclude that the series does not have a unit root if the statistic reported is less than –1.64 (at the 5% level). c and c, t indicate that a constant term and a constant term plus a trend components are included in the regression.
(3) $-\infty$ means that we obtain a very large negative value.

[1] The nominal money stock and the real GDP are expressed in US dollars.

Before we estimate our inflation models we need a measure of the long run demand for real money. To this end we estimate equation with several combinations of the explanatory variables and this yields three models (models 1 through 3, in Tables 2 and 3a, 3b, 3c below). We make use of estimators that exploit the information available in the cross-sections in order to obtain more precise estimates of the average parameters in the model. To this effect, we consider two types of panel estimators. Table 2 is made up of the results of long-run equations using a Fully Modified OLS (FMOLS) estimator. This estimator, proposed by Pedroni (2001), allows for heterogeneous slopes across the countries in addition to correcting for endogenous bias and serial correlation. Tables (3a) to (3c) also report results based on mean group estimators as proposed by Pesaran and Smith (1995) and Pesaran et al. (1999). This allows to make explicit the speed of adjustment to the long-run equilibrium.

From Table 2, our main findings are as follows. The estimates match prior expectations from theory. The specifications show a positive income elasticity close to unity, which is consonant with the quantitative theory hypothesis. An increase in real income thus results in a proportionate increase in the demand for real money. The demand for money is negatively related to the rate of inflation. The order of magnitude of the semi-elasticity is small (around -0.3). The explanation for this is that low inflation in the Euro area prevents strong substitution effects from monetary assets to real assets or foreign currencies. The demand for real money is positively affected by the short-term interest rate and negatively correlated with the long-term interest rate. However, the semi-elasticity for the former is not statistically significant. This is surprising, but it may illustrate the fact that when one uses a broad definition of money (M3 in our case) interest rates with the longest maturity better capture the effect of financial asset substitution. We also observe that the coefficient of the US bond yield is not statistically significant. The explanation of this can be found in the fact that, over the period under examination, the US interest rates have driven the European rates in a context of a 30-year trend of increasing integration across markets[1]. Thus, in the regression, the two rates are likely to be collinear. Eventually, we find the expected positive sign for the real effective exchange rate and this suggests a negative impact on money demand of a depreciating currency. This plays in favor of the currency substitution hypothesis.

Table 2. Pedroni FMOLS estimator – *t*-ratios in parentheses [(1)] Endogenous variable: real money demand

	Model 1	Model 2	Model 3
Real *GDP*	0.969* (81.78)	0.969* (81.17)	0.900* (58.05)
Inflation	-0.268* (-5.26)	-0.260* (-5.04)	-0.295* (-6.11)
Long-term interest rate	-0.008* (-8.17)	-0.0082* (-7.89)	-0.0074* (-7.24)
Short-term interest rate	0.00074 (1.82)	0.0007 (1.64)	0.00003 (0.84)
US bond yield	-	-0.0007 (0.09)	-
Real effective exchange rate	-	-	0.234* (6.25)
Panel cointegration tests[(2)]			
Panel *v*	2.54**	2.56**	1.85**
Panel ρ	-4.20**	-2.80**	-3.22**
Panel *PP*	-12.96**	-11.97**	-13.39**
Panel *ADF*	-8.27**	-7.53**	-6.72**
Group ρ	-3.35**	-1.95**	-2.21**
Group *PP*	-14.00**	-12.84**	-14.06**
Group *ADF*	-9.62**	-8.68**	-8.36**

Notes: (1) The FMOLS estimator is constructed by making corrections to the OLS estimator for the endogeneity of the regressors and serial correlation of the residuals. The endogeneity correction is achieved non-parametrically. To deal with the problem of spatial correlation, prior to the estimation, we first regress the individual series on yearly time dummies and work with the residuals of these regressions. The dummies are intended to capture shocks that are shared across the different members of the panel and thus to remove a potential common factor problem. An asterisk (*) implies significance at the 5% level. (2) All the statistics are distributed as standard normal asymptotically. The panel *v* rejects the null of no cointegration for large positive values (here for values higher than 1.64 at the 5% level) whereas the other six reject it with large negative values (here for values less than –1.64 at the 5% level). ** indicates the rejection of the null of no cointegration at the 5% level.

[1] For recent empirical studies concerning the synchronization of the US and the European interest rates, see Hammersland (2004) and Chinn and Frankel (2005).

As shown by the panel cointegration tests, a cointegration relationship between real money and the explanatory variables is accounted for. Indeed the Panel v statistic is well above 1.64 and the other six statistics lie under –1.64. However, equation assumes an instantaneous adjustment of the real money balance to its desired level. Such an instantaneous equilibrium state between the real money supply and the real demand for money is unlikely given the existence of transaction costs. We thus need to make a distinction between the long- and short-run behavior in the money market by specifying an error correction mechanism of actual real cash balances toward their desired (long-run) level. In this view, Tables (3a) through (3c) report some estimations based on the mean group (MGE) and pooled mean group (PMGE) estimators. The three tables correspond to the models 1 through 3 respectively.

Table 3a. Pooled mean group and mean group estimators – model 1 endogenous variable: real money demand

	PMGE(2)	MGE		
Long-run coefficients				
	Coefficients (*t*-ratios)	Coefficients (*t*-ratios)	*h*-test(1)	*P*-value
Real GDP	1.015* (50.519)	0.948* (14.165)	1.08	0.30
Inflation	-0.124* (-2.22)	-0.185 (-1.56)	0.34	0.56
Long-term interest rate	-0.005* (-4.00)	-0.009* (-2.12)	0.87	0.35
Short-term interest rate	0.003* (4.221)	-0.001 (-0.447)	2.44	0.12
Short-run coefficients(3)				
	Coefficients (*t*-ratios)	Coefficients (*t*-ratios)		
Real GDP	1.01* (45.90)	0.989* (10.08)		
Inflation	-0.124* (41.33)	-		
Long-term interest rate	-0.005* (-46.31)	-0.009* (-2.12)		
Short-term interest rate	0.003* (46.31)	-		
Δ real GDP(-1)	-0.073** (-1.762)	-		
Δ inflation	-0.082** (-1.744)	-		
Error-correction coefficients	-0.995* (-47.38)	-1.026* (-30.08)		

Notes: * statistically significant at the 5% level of significance and ** statistically significant at the 10% level of significance. (1) *h*-test: Hausman test of poolability. The test is constructed as a test of equivalence between the pooled mean group and the mean group estimates. Probability values are provided for this test: *p*-values larger than 0.05 indicate acceptance of the null of poolability. (2) The mean group estimates have been used as initial estimates of the long-run parameters for the PMGE. To deal with the problem of spatial correlation, prior to the estimation, we first regress the individual series on yearly time dummies and work with the residuals of these regressions. The dummies are intended to capture shocks that are shared across the different members of the panel and thus to remove a potential common factor problem. (3) We report the short-run coefficients that are statistically significant (either at the 10% or the 5% level). The Schwarz coefficient has been used to select the lag orders.

Table 3b. Pooled mean group and mean group estimators – model 2 endogenous variable: real money demand

	PMGE(2)	MGE		
Long-run coefficients				
	Coefficients (*t*-ratios)	Coefficients (*t*-ratios)	*h*-test(1)	*P*-value
Real GDP	0.991* (51.00)	0.922* (19.73)	2.65	0.10
Inflation	-0.226* (-4.17)	-0.210 (-1.703)	0.02	0.89
Long-term interest rate	-0.004* (-3.77)	-0.008** (-1.90)	0.74	0.39
Short-term interest rate	0.002* (3.091)	-0.001 (-0.597)	3.51	0.06
US bond yield	-0.000 (0.253)	-0.000 (-0.018)	0.41	0.52

Table 3b (cont.). Pooled mean group and mean group estimators – model 2 endogenous variable: real money demand

		PMGE[2]	MGE		
Short-run coefficients[3]					
		Coefficients (*t*-ratios)	Coefficients (*t*-ratios)		
Real GDP		0.969* (57.0)			
Inflation		-0.221* (-55.25)			
Long-term interest rate		-0.004* (-55.72)			
Short-term interest rate		0.002* (55.72)			
Δ real GDP(-1)		0.032** (1.783)			
intercept		-0.001* (-7.29)			
Error-correction coefficients		-0.977* (-54.27)	-0.98* (-24.11)		

Notes: see footnotes of Table 3a.

Table 3c. Pooled mean group and mean group estimators – model 3. Endogenous variable: real money demand

		PMGE[2]	MGE		
ong r n coe icients					
		Coefficients (*t*-ratios)	Coefficients (*t*-ratios)	-test[1]	*P*-value
Real GDP		0.949* (38.54)	0.80* (12.57)	6.45	0.01
Inflation		-0.232* (-4.37)	-0.359* (-3.109)	1.53	0.22
Long-term interest rate		-0.002 (-1.608)	-0.004 (-0.966)	0.30	0.58
Short-term interest rate		0.001** (1.78)	-0.003 (-1.25)	2.80	0.09
Real effective exchange rate		0.134* (2.66)	0.337* (2.697)	3.15	0.08
Short-run coefficients[3]					
		Coefficients (*t*-ratios)	Coefficients (*t*-ratios)		
Real GDP		0.915* (38.12)	0.817* (15.64)		
Inflation		-0.224* (-37.33)	-0.406* (-3.10)		
Long-term interest rate		-0.002* (-38.10)	-		
Short-term interest rate		0.001* (38.10)	-		
Real effective exchange rate		0.129* (43.00)	0.409* (2.36)		
Error-correction coefficients		-0.965* (38.6)	-1.061* (-20.71)		

Notes: see footnotes of Table 3a.

The PMGE method restricts the long-run coefficients to be equal across the countries (they are pooled), while the short-run coefficients are estimated individually and then averaged. The estimates are based on the application of the maximum likelihood approach and a Newton-Raphson algorithm to the following specification:

$$\Delta\left(m_{it} - p_{it}\right) = j_i \left(m_{it\text{-}1} - p_{it\text{-}1}\right) + \beta_i' X_{it} + \sum_{j=1}^{p\text{-}1} \lambda_{ij} \Delta\left(m_{it\text{-}j} - p_{it\text{-}j}\right) + \sum_{j=0}^{q\text{-}1} \mu_{ij} \Delta X_{it\text{-}j} + \eta_i + \nu_{it}. \tag{7}$$

For each country i, where ε_{it} is the disturbance term and X_{it} designates the vector of the explanatory variables. If the roots of this equation lie outside the unit circle, then the model is error-correcting ($\phi_i <$

0) and there exists a long-run relationship between the demand for real money and its determinants defined by:

$$m_{it} - p_{it} = -\left(\beta_i' / j_i\right) X_{it} + \varepsilon_{it}. \quad (8)$$

The long-run coefficients are equal to $\alpha_i = -\left(\beta_i' / \phi_i\right)$; the long-run homogeneity hypothesis is thus characterized by $a = a_i$ for every country i. Further, we calculate the mean group estimator which is an average of both the short- and long-run coefficients. We test for long-run homogeneity using a joint Hausman test based on the null hypothesis of equivalence between the PMGE and MGE estimator[1]. If the long-run coefficients are homogenous, then the MGE estimates are consistent and efficient (h-test in the tables). In the tables, the values of the h-test above 5% indicate acceptance of the null of poolability for the long-run coefficients. We can therefore consider that there are no significant statistical differences between the two estimators. The only exception is the coefficient of the real GDP (in model 3) for which the null of equivalent coefficients across the countries is rejected.

In all the regressions, the error-correction coefficients are very close to unity, implying that the short- and long-run coefficients are nearly equal. This suggests that the pressure on money demand to return to its long-run equilibrium is rather strong and that the adjustment time may be instantaneous. The short-term interest rate, which has no impact when we consider average-based long-run coefficients (the FMOLS or MGE estimator), does affect the demand for real money in the pooled-based regressions (PMGE). The corresponding coefficient is statistically significant either at the 10% or at the 5% level of significance with a semi-elasticity coefficient varying between 0.001 and 0.003. The US bond yield remains non significant. As for the other variables, we find that the real GDP positively affects real money demand, while both the inflation and the long-run interest rate have a negative influence (just as with the FMOLS estimator). Significant influences of the latter are found when we consider the PMGE estimator. Our preferred model is model 3 estimated with PMGE.

2.3. Empirical results for the inflation models. *2.3.1. Inflation model with the real money gap as a regressor.* Equation (5) is estimated as follows. To deal with the endogeneity problem, we use instruments and apply a GMM estimator (instead of a standard IV estimator) to gain efficiency by exploiting additional moments restrictions. The t-statistics are computed using heteroscedastic and serial correlation-consistent standard errors. To deal with the problem of spatial correlation, prior to the estimation, we first regress the individual series on yearly time dummies and work with the residuals of these regressions. These dummies are intended to capture shocks that are shared across the different members of the panel and thus to remove a potential common factor. Finally, to avoid colinearity problems, we consider lagged inflation and core inflation separately in the regressors (instead of the inflation-gap). The results are shown in Table 4 (in the table, the real money-gap is the estimated residual from the PMGE estimation of model 3 in Table 3b).

As is seen, the coefficient of the lagged real-money gap is positive and statistically significant. We can interpret this finding as an indication of the predictive power of the real money-gap for future inflation. We note that the coefficient of the output-gap is significant (at the 10% level of significance) and that the core inflation is also significant at the 5% level of significance.

Table 4. Estimates of the inflation model – GMM estimator with robust errors

	Inflation model with the real money-gap[(1)]	Inflation model with the nominal money growth gap[(2)]
Inflation (-1)	-0.10 (-0.67)	-0.05 (-0.12)
Output-gap (-1)	0.279** (1.74)	0.39** (1.77)
Core inflation (-1)	1.02* (6.55)	0.94* (2.42)
Real money-gap(-1)	0.057* (2.45)	-
Nominal money growth gap	-	0.02** (1.81)
Tests of the validity of the instruments: - Fisher statistics and *p*-values (in parentheses)		
Inflation	3.55 (0.007)	3.55 (0.007)
Output-gap	1.74 (0.024)	2.37 (0.0008)

[1] See Pesaran et al. (1996).

Table 4 (cont.). Estimates of the inflation model – GMM estimator with robust errors

Regressor	Inflation model with the real money-gap	Inflation model with the nominal money growth gap
Core inflation	0.63 (0.891)	0.69 (0.83)
Real money-gap	5.35 (0.000)	-
Nominal money growth gap	-	6.56 (0.000)
Sargan tests: over-identifying restrictions – Chi-squared statistics and -values		
	Inflation model with real money-gap	Inflation model with nominal money growth gap
	20.78 (0.236)	20.27 (0.260)

Note: *t*-ratio are in parentheses * statistically significant at the 5% significance level. ** statistically significant at the 10% significance level (1) List of instruments: inflation (lags 2 to 5), short-run real interest rate (lags 5 to 8), core inflation (lags 5 to 8), output-gap (lags 3 to 6), real money-gap (lag 7), inflation-gap (lags 8 to 11). (2) List of instruments: inflation (lags 2 to 5), short-run real interest rate (lags 5 to 8), core inflation (lags 5 to 8), output-gap (lags 5 to 8), nominal money growth gap (lag 8), inflation-gap (lags 8 to 11).

2.4. Inflation model with the nominal money growth gap as a regressor. To obtain a measure of the target value of the nominal money growth, Δm^*_{it}, we estimate equation (4) using estimators similar to those used to estimate the real money demand and find two estimated coefficients that are statistically significant with $\hat{\delta}_1$ = -1,13 and $\hat{\delta}_2$ = 2.83. Once this is done, we compute the static forecasts and estimate the following equation by GMM:

$$\pi_{it} = \beta_1 \pi_{it\text{-}1} + \beta_2 \left(y_{it\text{-}1} - y^*_{it\text{-}1} \right) + \beta_3 \left(\pi_{it\text{-}1} - \overline{\pi}_{it\text{-}1} \right) + \\ + \beta_4 \left(\Delta m_{it\text{-}1} - \Delta m^*_{it\text{-}1} \right) + \beta_5 \overline{\pi}_{it\text{-}1} + \eta_{it}. \quad (9)$$

The results are shown in Table 4. We see that the coefficient of the nominal money growth-gap is statistically significant. To check the robustness of our estimations, we test for the validity of the instruments used in the GMM regressions. For each exogenous variable in the inflation equation, we first test for the joint statistical significance of the instruments. As is seen in Table 4, except for the inflation target, the instruments can be considered as good predictors of the exogenous variables. Indeed, the *p*-values of the Fisher test are less than 5% and this induces the rejection of the null hypothesis of no correlation between the explanatory variables and the instruments. Further, we apply the Sargan test to see whether the instruments can be considered as exogenous. As the *p*-values of the tests are higher than 5%, we conclude that the instruments are independent of the error term in the inflation equations.

3. Out-of-sample forecasts

We now examine which of the two variables (the nominal money-growth rate or the real money-gap) is the most informative about future inflation. In this view, for purpose of illustration, we select countries according to the following criteria: some are member of the euro and the others are not: Austria, Denmark, France, Italy, the Netherlands, Norway, Sweden and the United Kingdom. We consider two types of horizons relevant for monetary policy: short-term and long-term horizons. As an illustration of short-run predictions, we compute one-quarter ahead forecasts. For the longer time horizons, we consider one-year and two-year ahead predictions[1]. The forecasts are computed over the period from 1999:1 through 2004:1. As in Gerlach and Svensson (2003) we assess the stability of the relationships prevailing until 1998:4 for the period after 1999:1. Accordingly, the forecasts are based on the estimates presented in the preceding sections and on the actual values of the explanatory variables. More specifically, the forecasts are obtained from the estimation of the following equations:

$$\pi_{it} = \beta_1 \pi_{it\text{-}j} + \beta_2 \left(y_{it\text{-}1} - y^*_{it\text{-}j} \right) + \beta_3 \left(\pi_{it\text{-}j} - \overline{\pi}_{it\text{-}j} \right) + \\ + \beta_4 \left(\hat{\varepsilon}_{it\text{-}j} \right) + \beta_5 \overline{\pi}_{it\text{-}j} + \eta_{it}, \quad (10)$$

and

$$\pi_{it} = \beta_1 \pi_{it\text{-}j} + \beta_2 \left(y_{it\text{-}j} - y^*_{it\text{-}j} \right) + \beta_3 \left(\pi_{it\text{-}j} - \overline{\pi}_{it\text{-}j} \right) \\ + \beta_4 \left(\Delta m_{it\text{-}j} - \Delta m^*_{it\text{-}j} \right) + \beta_5 \overline{\pi}_{it\text{-}j} + \eta_{it}. \quad (11)$$

Where j = 1, 4, 8. The estimated equations (11) and (12) are presented in Table 5.

In Table 6, we report the short-term horizon forecasts (one-quarter ahead forecasts). We compute the root mean squared errors of the predictions and discuss the predictive accuracy of the forecasts by examining a battery of test statistics. The numbers in bold correspond to the *p*-values of the cases where the inflation model with the nominal money growth rate yields forecasts that are better than those obtained with the real money-gap. As is seen, the predictions arising from the inflation equation with the nominal money

[1] One-year and two-year ahead predictions cannot be obtained from equations (5) and (10), consequently, these equations have to be re-specified and re-estimated equations (11) and (12).

growth rate are better than those coming from the equation with the real money-gap in the countries that gave up their national currency on January 1999 by adopting the Euro (France, Italy and the Netherlands)[1]. We obtain a similar conclusion for Denmark. The result concerning Denmark can be explained by the fact that the Danish central bank pursues a monetary policy that ensures a stable Krone vis-à-vis the Euro. In the other countries, the predictions arising from the inflation equation with the real money gap are quite similar to those obtained from the equation with the nominal money growth. Considering long-term forecasts[2] (Table 7), the conclusions seem to be the same: the nominal money growth is still determinant in three out of four countries that use the Euro (Austria, France and the Netherlands).

Table 5. Estimates of the inflation model – GMM estimator with robust errors 1-year and 2-year-ahead inflation against the current values of the explanatory variables

	Inflation model with the real money-gap		Inflation model with the nominal money growth gap(2)	
	1-year(1) (j = 4)	2-years(2) (j = 8)	1-year(3) (j = 4)	2-years(4) (j = 8)
Inflation (-j)	-0.056 (-0.48)	0.24* (2.09)	-0.18 (-1.52)	0.03 (0.60)
Core inflation (-j)	0.857* (2.69)	0.21 (1.50)	0.95* (7.17)	0.57* (10.79)
Output-gap (-j)	0.827* (6.81)	0.309* (2.74)	1.11* (3.88)	0.17 (0.59)
Real money-gap(-j)	0.351* (3.90)	0.05** (1.88)	-	-
Nominal money growth gap(-j)	-	-	0.003 (0.194)	0.04* (3.54)
Tests of the validity of the instruments: - Fisher statistics and p-values (in parentheses)				
Regressor	Inflation model with the real money-gap		Inflation model with the nominal money growth gap	
Inflation	2.93 (0.0024)	4.02 (0.0002)	2.81 (0.007)	3.53 (0.0007)
Output-gap	1.54 (0.042)	1.78 (0.035)	1.80 (0.038)	2.06 (0.03)
Core inflation	1.17 (0.258)	1.36 (0.10)	1.19 (0.275)	1.63 (0.03)
Real money-gap	6.12 (0.00)	3.16 (0.00007)	-	-
Nominal money growth gap	-	-	3.57 (0.00002)	3.01 (0.00002)
Sargan tests: over-identifying restrictions – Chi-squared statistics and p-values				
	Inflation model with the real money-gap		Inflation model with the nominal money growth gap	
	33.40 (0.121)	38.14 (0.09)	15.58 (0.11)	27.67 (0.07)

Note: t-ratios are in parentheses * statistically significant at the 5% significance level. ** statistically significant at the 10% significance level. (1) List of instruments: inflation (lags 5 to 12), core inflation (lags 5 to 12), output-gap (lags 5 to 12), real money-gap (lags 5 to 12). (2) List of instruments: inflation (lags 9 to 16), core inflation (lags 9 to 16), output-gap (lags 9 to 16), real money-gap (lags 9 to 16). (3) List of instruments: inflation (lags 6 to 12), nominal money growth gap (lags 6 to 12). (4) List of instruments: inflation (lags 10 to 17), core inflation (lags 10 to 16), output-gap (lag 10), nominal money growth gap (lags 10 to 16).

Table 6. One-quarter ahead forecasts: RMSE and predictive accuracy tests (p-values in brackets) Inflation model with real money-gap versus inflation model with nominal money growth

	Austria	Denmark	France	Italy	Neth	Norway
RMSE1(1)	0.80×10^{-2}	0.66×10^{-2}	0.16×10^{-1}	0.45×10^{-2}	0.156×10^{-1}	0.29×10^{-1}
RMSE2(1)	0.64×10^{-2}	0.72×10^{-2}	0.98×10^{-2}	0.48×10^{-2}	0.10×10^{-1}	0.28×10^{-1}
Austria	1.196 (0.231)	0.00 (1.00)	0.892 (0.372)	0.642 (0.846)	-1.253 (0.223)	-1.143 (0.253)
Denmark	-1.929 (0.053)	-1.876 (0.06)	-2.220 (0.026)	1.202 (0.332)	1.799 (0.085)	1.155 (0.247)
France	7.348 (0.00)	1.876 (0.06)	2.463 (0.014)	0.369 (0.989)	-4.111 (0.0005)	-2.277 (0.022)
Italy	-3.373 (0.0007)	-2.293 (0.022)	-2.281 (0.022)	1.175 (0.350)	1.829 (0.081)	1.413 (0.157)

[1] A different conclusion is obtained for Austria.
[2] We only report the results concerning the two-year horizon forecast, since those concerning the one-year forecasts were very similar.

Table 6 (cont.). One-quarter ahead forecasts: RMSE and predictive accuracy tests (p-values in brackets) Inflation model with real money-gap versus inflation model with nominal money growth

	Austria	Denmark	France	Italy	Neth	Norway
Netherlands	8.484 (0.00)	1.042 (0.297)	2.129 (0.03)	0.429 (0.976)	-3.822 (0.0009)	-2.201 (0.027)
Norway	1.106 (0.268)	0.208 (0.834)	0.821 (0.411)	0.951 (0.547)	-1.919 (0.068)	-1.186 (0.235)
Sweden	1.319 (0.1869)	-1.459 (0.144)	-0.912 (0.361)	0.961 (0.537)	-1.043 (0.308)	-1.217 (0.223)
UK	-0.871 (0.383)	-0.208 (0.834)	-0.152 (0.879)	1.014 (0.486)	0.407 (0.687)	0.433 (0.665)

Note: (1) RMSE1 and RMSE2 are the root mean squared error of the predictions based respectively on equation (8) or (11) and equation (10) or (12). (2) The different columns are: AS: asymptotic test, SI: sign test, WI: Wilcoxon's test, NB: naive benchmark test, MGN: Morgan-Granger-Newbold's test, MR: Meese-Rogoff's test. The null hypothesis is the hypothesis of equal accuracy of the predictions. The loss function is quadratic. The test statistics follow asymptotically different distributions: N (0,1) for the asymptotic test, the sign test, the Wilcoxon's test, the Meese-Rogoff's test, F (T_0, T_0) for the naive benchmark tests and a $t_{T0\text{-}1}$ for the Morgan-Granger-Newbold's test (where T_0 is the number of predicted observations). The Meese-Rogoff test statistic is computed with the Diebold-Rudebusch covariance matrix estimator. The truncation lag is 10 for the asymptotic test and is given by the integer part of $T_0^{4/5}$ for the Meese-Rogoff's test. For a detailed presentation of the different tests, the reader can refer to Diebold and Mariano (1995).

Table 7. Two-year-ahead forecasts: RMSE and predictive accuracy tests (p-values in brackets). Inflation model with real money-gap versus inflation model with nominal money growth

	Austria	Denmark	France	Italy	Neth	Norway
RMSE1(1)	0.11×10^{-1}	0.82×10^{-2}	0.16×10^{-1}	0.55×10^{-2}	0.19×10^{-1}	0.33×10^{-1}
RMSE2(1)	0.53×10^{-2}	0.71×10^{-2}	0.10×10^{-1}	0.47×10^{-2}	0.12×10^{-1}	0.33×10^{-1}
	AS(2)	SI(2)	WI(2)	NB(2)	MGN(2)	MR(2)
Austria	2.74 (0.006)	2.324 (0.020)	2.55 (0.01)	0.211 (0.997)	-3.55 (0.003)	-2.08 (0.037)
Denmark	1.658 (0.097)	0.00 (1.00)	0.258 (0.796)	0.752 (0.712)	-1.752 (0.100)	-1.613 (0.106)
France	3.809 (0.0001)	2.00 (0.034)	2.12 (0.034)	0.435 (0.946)	-2.41 (0.029)	-1.69 (0.09)
Italy	0.840 (0.401)	1.00 (0.317)	1.499 (0.133)	0.741 (0.721)	-1.392 (0.183)	-1.238 (0.215)
Netherlands	5.499 (0.000)	2.50 (0.012)	3.05 (0.002)	0.381 (0.968)	-5.167 (0.0001)	-2.29 (0.02)
Norway	0.849 (0.395)	-1.50 (0.133)	-0.258 (0.796)	0.982 (0.513)	-0.253 (0.803)	-0.315 (0.752)
Sweden	1.653 (0.098)	-0.50 (0.617)	0.103 (0.917)	0.909 (0.574)	-1.152 (0.267)	-0.856 (0.392)
UK	2.377 (0.017)	1.00 (0.317)	1.861 (0.062)	0.820 (0.651)	-1.908 (0.075)	-1.407 (0.159)

Note: see footnote Table 6.

Conclusions

This paper is an empirical contribution to the ongoing debate on the relative performance of a nominal money growth based indicator or the real money-gap in predicting future inflation. Our conclusion is that the nominal money-growth indicator has a higher predictive power of future inflation in the countries that adopted the Euro and where the monetary policy strategy is based on the one announced by the European Central Bank. In the other countries, monetary aggregates other than the nominal, for instance the real money gap, are equally informative for future inflation.

Implications

The fact that, in the Euro countries, nominal money growth dominates the real money gap in predicting inflation provides empirical support for the policies adopted by the European Central Bank since the adoption of the Euro. Indeed, the Eurosystem's monetary policy consists in maintaining price stability and choosing a nominal money growth indicator to achieve this objective (see European Central Bank, 1999).

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