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Money demand stability, monetary overhang and inflation forecast in the CEE countries

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Abstract

This paper first shows that the long-run money demand in Central and Eastern European (CEE) countries is better described by an open-economy model (OEM), which considers a currency substitution effect, than by a closed-economy model (CEM) used in several previous studies. Second, from the estimated models we derive two different measures of monetary overhang. Then we compare the ability of the OEM-based and the CEM-based measures of monetary overhang to predict inflation in the CEE countries, namely the Czech Republic, Hungary and Poland. While we cannot detect a significant difference of forecast accuracy between the two competing models, we show that the OEM-based forecast model that reveals a stable long-run money demand encompasses the CEM-based version for the CEE countries.

Keywords: money demand stability; monetary overhang; inflation forecasts; currency substitution; CEE countries

JEL codes: E41, E47, E52, F41

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1. Introduction

The money demand function, one of the most investigated macroeconomic relationships, is meant to help the monetary authority to understand what motivates economic agents to hold money. Therefore, the monetary authority can decide which monetary targets and policies are recommended under specific economic conditions. The stability of money demand function states that the money supply has a potential impact on both economic activity and inflation. Otherwise said, a stable money demand shows how effective the use of monetary aggregates is, in the conduct of monetary policy.

Money demand stability is derived from the quantity theory of money, where money supply is exogenous, and money supply changes pass-through production and inflation. At the same time, the velocity of money is supposed to be stable. Given the fact that velocity represents a linear combination of money, production and inflation, testing the stability of velocity is equivalent with testing the stability of money demand (Bahmani, 2008). Although the ‘modern monetary theory’ and the New-Keynesian models generally omit the money demand shocks in the empirical analysis of monetary policy (Biscarri et al., 2010), noteworthy recent studies (e.g. Teles et al., 2015) conclude that “quantity theory is still alive.”

A glance through the literature reveals that the investigation of money demand stability has multiple implications. First, it helps to make an adequate choice of a monetary policy instrument because the instability of money demand is a major determinant of liquidity preference (Kumar et al., 2013). Second, a stable money demand implies a stable money multiplier, ensuring a correct predication of the effects generated by money-supply shocks on the aggregate income, according to the monetarist view (Narayan, 2010). Third, and very important, the stability of money demand provides valuable information for the nexus between money and inflation.

For a long time, economists investigated the causes of high-inflation episodes, and one of the possible explanations is provided by the demand for money (Eckstein and Leiderman, 1992). On the one hand, a stable money demand function is associated with a long-run positive correlation between the money in circulation and the level of inflation, showing that money can predict inflation. A key issue here is the identification of the ‘monetary overhang’ or excess liquidity, because movements in money supply do not always match movements in money demand (El-Shagi and Giesen, 2013). On the other hand, there is an increased interest in estimating the welfare cost of inflation (Attanasio et al., 2002). The calculation of the welfare

cost assumes that the private sector expects the current inflation rate, with positive effects on income and on inflation-targeting strategy (Miller et al., 2017).

Our paper adds to the existing menu of studies that investigate the stability of money demand and the role of monetary overhang in predicting inflation in the Central and Eastern European (CEE) countries. The evaluation of money demand stability remains a subject of interest for these countries because their monetary policy tends to be oriented toward discretion rather than rules (Czirák and Gillman, 2006). After a successful disinflation process recorded by these countries, their Euro Area accession became a subject of interest in the context of a new European Union (EU) framework and developing strategy. From the perspective of the Euro Area enlargement, a stable money demand function creates a good pre-condition for the euro adoption (Fidrmuc, 2009).¹ Further, these countries adopted inflation-targeting regimes. Therefore, investigating the stability of money demand shows to what extent unexpected monetary shocks challenge the performance of inflation-targeting strategies.

Despite the intensive empirical efforts to study the stability of money demand in CEE countries, there is no consensus regarding the existence, or the absence of stability. This casts some doubt about the role of money in forecasting inflation. The mixed empirical findings might be explained by a different specification of money demand function used by previous studies. Cagan's (1956) constant semi-interest elasticity model serves as starting point for many empirical works on money demand, which focus on the interest rate role and investigate the stability of money demand functions. Several early studies (e.g. Budina et al., 1995; Narayn, 2010) employ this closed-economy formulation to compute the stability of money demand in CEE countries. Other previous works (e.g. Dreger et al., 2007; Fidrmuc, 2009) resort to a general formulation of the money demand function as in Leventakis (1993), and consider open-economy factors.

Different from the existing empirical literature, we test the stability of money demand by comparing two competing money demand models, namely the classic Cagan's (1956) closed-economy model (CEM), and the Albulescu et al.'s (2017) open-economy model (OEM). This comparison allows us to explain the mixed findings that are reported in the existing literature for the stability of money demand. In addition, the OEM model we use is a micro-founded

¹ The White Paper on the future of Europe released by the European Commission in March 2017, after the Brexit decision, designs several scenarios for the EU27 until 2025. Given the new commitments for a stronger EU and the election results in the old EU members with a strong Euroscepticism current (i.e., France and Netherlands), the last scenario, namely "Doing much more together" seems to be plausible. This means that new member states are encouraged to join the Euro Area while a stronger fiscal coordination is necessary. In this context, understanding the determinants of money demand for the CEE countries is crucial.

model, which, different from Leventakis's (1993) specification, does not assume *ex-ante* the existence of a direct currency substitution between domestic currencies and the euro. This model is compatible with both currency substitution and currency complementarity effects, and is well adapted for the CEE countries, as it makes the assumption that the euro offers liquidity services to the domestic representative agent, while the reverse is supposed not to be true (for details, please refer to Albulescu et al., 2017). We resort to the Hansen's parameter instability test (Hansen, 1992) to assess the cointegrating relationship, and thus, the stability of money demand. Further, we estimate the long-run money demand parameters with the Fully Modified Ordinary Least Squares (FMOLS) method of Phillips and Hansen (1990).

Another contribution of our paper to the existing literature resides in the identification of the monetary overhang, starting from these two OEM and CEM long-run equations.² The purpose is to see to what extent the monetary overhang represents a good predictor of inflation in the Czech Republic, Hungary and Poland, over the period 1999 to 2016 (monthly data). For this purpose, we start from Stock and Watson's (1999) inflation forecasting equation (out-of-sample and combined forecasting approach). We further apply tests of equal forecast accuracy and encompassing tests (following Harvey et al., 1997, 1998). For robustness purpose, we reduce our sample to the period 2004 to 2016, to see if the monetary overhang represents a reliable predictor of inflation in CEE countries after their EU accession.

Our focus is on the 'advanced' CEE countries for two reasons. On the one hand, we are constrained by the data availability. As Cziráky and Gillman (2006) notice, evaluating the stability of money demand in CEE countries is problematic, because of the lack of confidence in data quality. In order to overcome this issue, we resort to international statistics for harmonization purpose, without considering national databases as most of the previous studies do. On the other hand, the selected countries have in place floating exchange rate regimes, and their degree of integration with the Euro Area is considered to be higher compared to other candidate countries (i.e., Bulgaria, Croatia or Romania).

We discover that in the long run, the open-economy specification of the money demand model gives more consistent results than the closed-economy version. Further, we show that the OEM-based monetary overhang is the best predictor of inflation for Hungary and Poland, especially at a long horizon. These findings are confirmed by our robustness analysis.

² The monetary overhang is associated with the residuals of the cointegration relationships. A positive monetary overhang represents a sign of inflationary pressure. If, for example, the money demand proves to be unstable, this will bias the inflation forecasting results (see Carstensen et al., 2009).

The rest of the paper is as follows. Section 2 presents the money stabilization literature. Section 3 is dedicated to the analysis of the demand for money stability, comparing the performance of a closed- and an open-economy money demand model. Section 4 addresses the monetary overhang and the inflation forecast in the selected CEE countries. Section 5 presents the robustness analysis. The last section concludes and presents the policy implications of our findings.

2. Money demand stabilization: a review of the literature

For a long time, the quantity theorist maintained that the demand for money is highly stable, and underlined its importance for transmission of monetary impulses into prices. However, during the 1980s, several elements, such as financial innovations and the deregulation process, put into question the concept of money demand stability (Lucas and Nicolini, 2015). Starting with Friedman and Kuttner (1992), who report a break in the cointegration relationship around 1980 for the United States (US), a new instability literature emerged, characterized by the consideration within the money demand function of substitutes for money. Further, the focus on monetary policy rules (Taylor, 1999) introduces the perception of policy irrelevance of money demand theory (Czirák and Gillman, 2006), although Alvarez et al. (2001) underline the ongoing role of money for the equivalence between interest rate rules and money-supply rules. Nevertheless, the role of money in forecasting inflation still is of great interest (Horváth et al., 2011).

The empirical literature on money demand stabilization usually resorts to simultaneous equations, cointegration analyses, and they apply consecrated stability tests, as recursive residuals approaches (CUSUM, CUSUMSQ), or Chow tests. Most of these studies address the case of developed economies, while few empirical works are oriented towards emerging economies and CEE countries.³

Early studies in this area focus on the US economy and show mixed evidence. Laumas and Mehra (1977) examine the stability of the US money demand using annual data for the period 1900-1974. Their varying parameter technique, which regresses money on income and

³ Estimates of the demand for money stability are familiar for developed countries (i.e., Ewing and Payne, 1999; Hondroyannis et al., 2001; Kumar and Webber, 2013; Jawadi and Sousa, 2013; Makin et al., 2017), and become of great interest for emerging economies. In the latter case, the stability of money demand and the monetary overhang are investigated *inter-alia* by Price and Nasim (1999) for Pakistan, by Wu et al. (2003) for Taiwan, and by Rao and Kumar (2009) for a set of Asian countries. Most of these works report stable money demand functions.

interest rates, shows that money demand is not stable. Opposite findings are advanced by Lin and Oh (1984), who employ switch-regression techniques and document money demand stability. These results are confirmed by recent works on the post-war stability of money demand in the US. Along this line, Arize et al. (2012) examine the long-run stability of money demand using the Johansen cointegration technique. Further, Miller et al. (2017) resort to nonlinear cointegration techniques and document the stability of money demand in the US.

The studies on the money demand stability in the EU are generally oriented on the Euro Area monetary stabilization as a whole, while few of them approach the case of Euro Area individual countries (for a recent review of the literature, please refer to Liu and Kool, 2018). Golinelli and Pastorello (2002) show that the area-wide money demand is more stable than the single-country one, while Brand and Cassola (2004) and Coenen and Vega (2001) reaffirm the stability of money demand at the aggregate level. Belke and Czudaj (2010) investigate the money demand at the aggregate level, and compare cointegrated VAR and single-equation techniques. The authors posit that the recent financial crisis has no noticeable impact on the stability of money demand. With a focus on the same aggregate level, De Santis et al. (2013) adopt a different approach and argue that a stable broad money demand for the Euro Area can be obtained by modelling cross-border international portfolio allocation. Further, applying the Johansen's (1995) cointegration method and the Nyblom's (1989) time-invariance parameter test, the authors report a strong co-movement between net cross-border portfolio flows and M3 velocity growth. Dreger and Wolters (2014) analyze the stability of money demand in the Euro Area and the inflation forecasting performances of a broad monetary aggregate. They state that the evolution of M3 is in line with the money demand.

Adopting a different strategy, Setzer and Wolff (2013) focus on the stability of money demand for the selected Euro Area countries, using disaggregated data. Their cointegration relationship is generated from a micro-founded money demand model, and the results show that the income and the interest rate elasticity remain stable over time. Capasso and Napolitano (2012) add another piece of evidence to the stability of money demand in the Euro Area countries. With a focus on Italy over the period 1977 to 2007, and using bounds-testing cointegration, the authors show that the introduction of the euro contributed to the money demand stability.

The interest for assessing the stability of money demand in the Euro Area increased after the recent global-crisis outburst, but it progressively vanished with the conduct of unconventional monetary policy by the European Central Bank (ECB). Similarly, the uncertainty generated by the crisis regarding the Euro Area enlargement, diminished the interest

for assessing the stability of money demand in CEE. However, studying the stability of money demand in the Euro Area candidate countries still remains a subject of great interest for researchers and policy makers.

The bulk of empirical literature on the monetary stabilization in the CEE countries is divided into panel data and time-series analyses. Given the fact that during the 1990s, the CEE transition economies registered noteworthy structural changes that make it difficult to access the data for a long sample period, a first set of studies uses panel data investigations. In this line, Dreger et al. (2007) employ panel cointegration methods and quarterly data for the period 1995 to 2004 and report the existence of a long-run relationship in the money demand equation. Similarly, Fidrmuc (2009) does not find any structural breaks in the series and concludes that the money demand is stable in CEE countries, even if it is largely influenced by the Euro Area interest rates and by the exchange rate against the euro, which might represent signs of instability. Applying the Hansen's (1992) parameter stability test for a panel of eight transitional economies, for the period 1995:01 to 2005:03, Narayan (2010) finds the existence of a long-run relationship, but reveals more cases of unstable money demand functions.

At the same time, starting with Chawluk and Cross (1997), the researchers become interested by the existence and the size of a monetary overhang to predict the inflation in the CEE countries. However, the first study that investigates the stability of money demand in individual post-communist economies is that of Buch (2001). Using a cointegration and error-correction framework and a CUSUM test for the coefficient stability, the author documents the stability of money demand in Hungary and Poland. Cziráky and Gillman (2006) estimate the money demand in Croatia using monthly data from 1994 to 2002. They find evidence for a stable money demand function, which represents the basis for inflation rate forecasting in Croatia. Subsequent studies enlarge the group of the analyzed countries. Applying a bounds-testing approach to error-correction modelling and cointegration, Bahmani and Kutan (2010) show that money demand in the case of the CEE countries is quite stable. Further, Bahmani-Oskooee et al. (2013) introduce uncertainty and monetary volatility in the money demand equation for six CEE and four other emerging economies. Their Pesaran et al.'s (2001) bounds-testing approach shows that the money demand is correctly specified and stable.

Nevertheless, to the best of our knowledge, no previous study investigates the stability of money demand that compares concurrent money demand functions (relying on closed- and open-economy models). Moreover, no previous money demand function employed for testing the monetary stabilization in the CEE countries is specifically designed to fit the particularities of the CEE economies. To fill in this gap, we compare the stability of money demand, resorting

to a closed- and an open-economy money demand model. We also compare the performance of the monetary overhang obtained, based on these competing models, in forecasting inflation in selected CEE countries.

3. Money demand in the long run

In this section, we describe two competing long-run money demand functions and present the estimated models for Hungary, Czech Republic and Poland. The monthly data period is 1999-01 to 2016-11. Details are given in Appendix A.

The first model is a standard closed-economy model that relates the real money demand to a scale variable (the real industrial production) and to a domestic interest rate. We employ the semi-log form used by Cagan (1956) and recommended by Ireland (2009), which explains the log of the real money demand by the log of the scale variable and the level of the interest rate:

$$\ln m_t = \alpha_0 + \alpha_1 \ln y_t + \alpha_2 r_t + \varepsilon_t \quad (01)$$

with $m_t = M_t/P_t$, where M, P, y and r are defined as money (M3 aggregate), prices, output (real industrial production) and domestic interest rate.⁴ α_0 is a positive constant, α_1 is the output elasticity and α_2 is the absolute value of the interest semi-elasticity of money demand.

The second model is a modified version of the open-economy model of Albuлесcu et al. (2017).⁵ In this model, the agents of CEE countries hold domestic and foreign assets (bonds and currencies). The foreign currency is the euro, which is supposed to offer liquidity services to the agents of the CEE countries. Currency substitution is demonstrated to relate the money demand to the interest rate spread between the CEE countries and the Euro Area (in addition to the scale variable and to the domestic interest rate). For this reason, we include the interest rate spread in the open-economy formulation of the money demand:

$$\ln m_t = \alpha_0 + \alpha_1 \ln y_t + \alpha_2 r_t + \alpha_3 (r_t - r_t^*) + \varepsilon_t \quad (02)$$

⁴ The M3 aggregate is represented by the broad money index (2010=100) and comes from the OECD database. Bruggeman et al. (2003), Carstensen et al., (2009) and Dreger and Wolters, (2014) use the M3 aggregate to check the stability of money demand in the Euro Area. The level of prices (consumer price index – 2010=100) and the real industrial production index (2010=100) comes from IMF-International Financial Statistics database. The interest rate (1-month money market rate) comes from the Eurostat database. Except for the interest rate, all the other series are seasonally adjusted as in Fidrmuc (2009).

⁵ To make relevant the comparison between CEM and OEM, we consider a semi-log form of OEM, whereas Albuлесcu et al. (2017) use a log-log specification. The semi-log form is also recommended by the fact that, during the analyzed time span, we have recorded periods with negative interest rates.

where r^* is the Euro Area interest rate. The parameter α_3 can be of any sign, depending on the elasticity of substitution between consumption and liquidity (which is a function of domestic and foreign money), and the elasticity of substitution between domestic and foreign money. The parameter α_3 is positive if substitutability between consumption and liquidity is higher than substitutability between foreign and domestic money, and negative otherwise (Albulescu et al., 2017). Assuming that liquidity and consumption are complements in the CEE countries (see, for example, the result of Albulescu et al., 2017), the ‘positivity’ of α_3 is a sign of complementarity between the CEE currencies and the euro.

In the long run, the model is supposed to perfectly fit the data so that the error term ε_t is zero. But in the short run, money demand deviates from the long-run function, giving rise to a nonzero stationary error term ε_t . The closed-economy version of the model is simply obtained with the restriction $\alpha_3 = 0$.

For each model and each country, we estimate the long-run money demand parameters with the FMOLS method of Phillips and Hansen (1990),⁶ and perform a cointegration test with the Hansen’s parameter instability test (Hansen, 1992).⁷ As the theoretical money demand model adds the restrictions that the output elasticity is one: $\alpha_1 = 1$, and that the interest-rate semi-elasticity is negative: $\alpha_2 < 0$, we also pay attention to these hypotheses.

Table 1 presents the long-run estimated parameters. It appears that the open-economy model offers more consistent results than the closed-economy model. Indeed, the estimated interest semi-elasticity is always negative and significant, whereas the same estimated parameter is positive in two out of three cases when the closed-economy specification is considered. Moreover, all the parameters of the open-economy model are significant, notably the interest spread between the CEE countries and the Euro Area. The estimated parameter α_3 is positive for the three CEE countries, which demonstrates that preference parameters are quite homogeneous among the CEE countries.

⁶ DOLS estimator of Saikkonen (1991) and Stock and Watson (1993) is another method used for estimating cointegration regressions. Both methods have similar asymptotic properties, but FMOLS is more frequently used to estimate money demand equations. For small samples such as ours, FMOLS non-parametrically presents the advantage that corrects for endogeneity and serial correlation to the OLS estimator.

⁷ Narayan (2010) used a similar approach in his panel data analysis for the stability of money demand in eight CEE countries.

Table 1. Estimation results

	Hungary		Czech Republic		Poland	
	OEM	CEM	OEM	CEM	OEM	CEM
α_0	-5.195*** (-10.764)	-4.867*** (-8.824)	-6.125*** (-9.310)	-3.012*** (-5.394)	-6.311*** (-20.310)	-7.007*** (-17.099)
α_1	1.093*** (11.083)	1.030*** (9.126)	1.319*** (9.439)	1.129*** (5.484)	1.367*** (20.961)	1.510*** (17.479)
α_2	-0.013* (-1.735)	0.002 (0.391)	-0.021* (-1.737)	-0.018 (-0.977)	-0.025*** (-2.962)	0.012*** (2.649741)
α_3	0.027*** (3.012)		0.096*** (6.098)		0.041*** (4.742)	

Notes: (i) The t-statistics are in bracket below the estimated long run parameters; (ii) ***, **, * means significance at 1%, 5% and 10% significance level.

It appears from Table 1, that the estimated output elasticities are not too far from the unity. For the OEM, the sign of α_2 is negative, meaning that a higher opportunity cost of holding the domestic and foreign money determines a decrease of money demand. In addition, the sign of α_3 is positive (and the coefficient $\alpha_3 > |\alpha_2|$), showing that a decrease in the opportunity cost of foreign money leads to an increase in domestic money demand. This result underlines the fact that CEE currencies and the euro are complements rather than substitutes, a result in line with the findings advanced by Albulescu et al. (2017).

Table 2 presents the results of the test of the hypothesis $\alpha_1 = 1$ for each country and each model. The hypothesis of a unitary output elasticity cannot be rejected for Hungary and the Czech Republic, whereas the hypothesis is rejected for Poland. Nevertheless, the estimated parameter is rather close to the unity, in particular in the case of the open-economy model.

Table 2. Tests of the hypothesis $\alpha_1 = 1$

	Hungary		Czech Republic		Poland	
	OEM	CEM	OEM	CEM	OEM	CEM
t-statistic (probability)	0.951 (0.342)	0.270 (0.787)	2.284 (0.023)	0.627 (0.530)	5.633 (0.000)	5.909 (0.000)

Table 3 presents the results of the cointegration tests performed according to the Hansen's parameter instability test (Hansen, 1992). The hypothesis that m, y and r are cointegrated is not plausible as the p-value of the cointegration test of the closed-economy model never exceeds 0.01. On the contrary, the hypothesis that m, y, r and r-r* are cointegrated is plausible, except for the Czech Republic where the assumption of stationarity of the error term ε_t is questionable.

To summarize, the open-economy specification of the money demand model gives more consistent results than the closed-economy version, consistent with a long-run relationship

between the variables for Hungary and Poland. The hypothesis of cointegration is more doubtful in the case of the Czech Republic.⁸

Table 3. Results of the cointegration tests (Hansen's parameter instability tests)

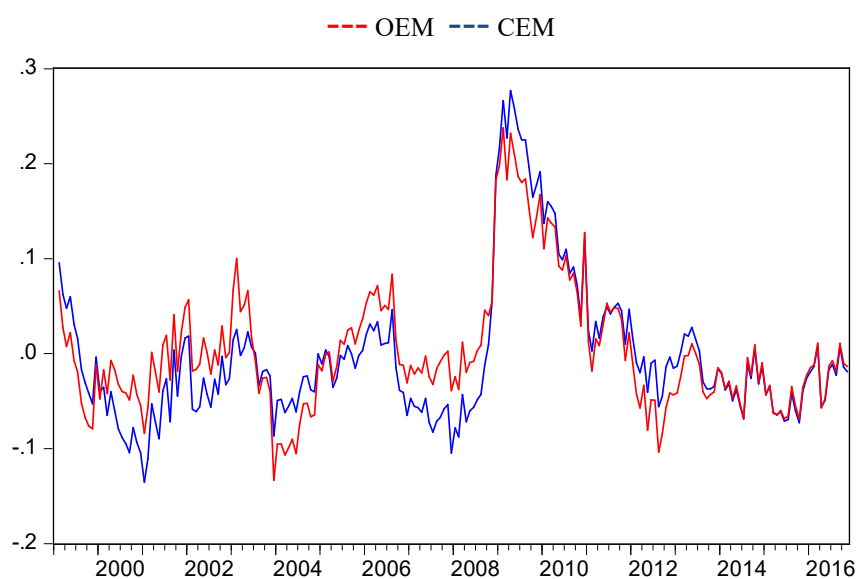
	Hungary		Czech Republic		Poland	
	OEM	CEM	OEM	CEM	OEM	CEM
Lc statistic (probability)	1.035 (0.019)	0.698 (0.010)	1.198 (< 0.01)	1.372 (< 0.01)	0.637 (0.123)	0.882 (< 0.01)

Next, we define the monetary overhang for each country and each model as the residual of the long-run money demand equations:

$$\hat{\varepsilon}_t = \ln m_t - \ln \hat{m}_t \quad (03)$$

Figures 1, 2 and 3 illustrate the differences between the two measures of monetary overhang.

Figure 1. Monetary overhang of Hungary



There are noticeable differences between these measures, especially for the Czech Republic and Poland. Moreover, we notice that the monetary overhang estimated from the OEM surpasses the monetary overhang estimated from the CEM for the three CEE countries during the pre-crisis period 2006-2008. We can also see that both measures of monetary overhang rocket at the beginning of the 2008 financial crisis.

⁸ To check the robustness of these results, we perform a similar analysis for the period 2004-2016 (the moment of the three CEE countries EU accession up to present). This set of results is presented in Appendix B and confirm the existence of a long-run relationship (see Tables B1, B2 and B3).

Figure 2. Monetary overhang of the Czech Republic

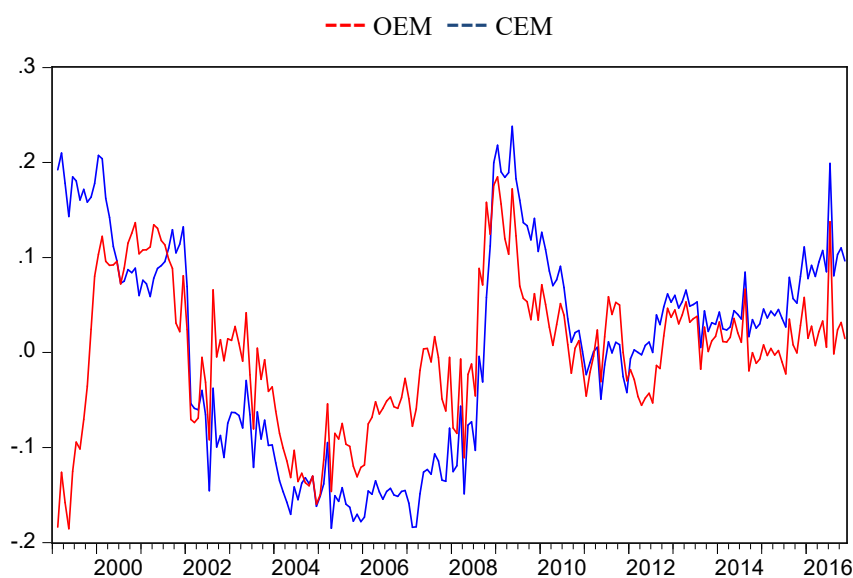
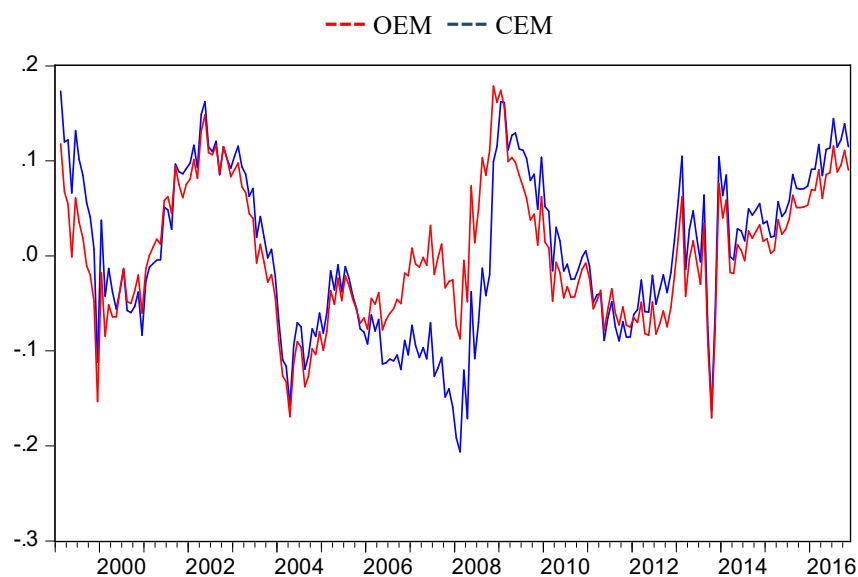


Figure 3. Monetary overhang of Poland



4. Monetary overhang and inflation forecast in the CEE countries

It is generally accepted that the existence of a stable long-run money demand function implies that the monetary overhang is a good leading indicator of inflation. Thus, we study and compare the inflation forecasting properties of the two measures of monetary overhang. For this purpose, we consider inflation forecasting for different forecast horizons h : a month ($h =$

1), a quarter ($h = 4$), a year ($h = 12$) and two years ($h = 24$).⁹ We follow Stock and Watson (1999) to specify the inflation forecasting equation:

$$\pi_{t+h}^h - \pi_t = \beta + \beta_\pi(L)\Delta\pi_t + \beta_{\hat{\varepsilon}}(L)\hat{\varepsilon}_t + \eta_{t+h} \quad (04)$$

where $\pi_t = 12 \times \ln(P_t/P_{t-1})$ is the annualized monthly inflation rate in the price level P_t , $\pi_{t+h}^h = (12/h) \times \ln(P_{t+h}/P_t)$ is the annualized h-period inflation rate, $\hat{\varepsilon}_t$ is one of the two measures of monetary overhang, Δ is the difference operator, $\beta_\pi(L)$ and $\beta_{\hat{\varepsilon}}(L)$ are lag polynomials, and η_{t+h} is an error term.

It is important to note that the CEM-based inflation forecasting equation is not nested in the OEM-based version (although the CEM is nested in the OEM). We have shown in the previous section that the OEM better describes the long-run money demand in CEE countries than the CEM does. Nevertheless, this result does not imply that the OEM-based monetary overhang has better forecasting properties for the future inflation than the CEM-based monetary overhang. To compare the forecasting performances of these two measures of monetary overhang, we implement tests of equal forecast accuracy and encompassing tests.

To estimate equation (04), we follow Carstensen et al. (2009) and fix the lag order of the polynomials $\beta_\pi(L)$ and $\beta_{\hat{\varepsilon}}(L)$ to 12, because the data are monthly.¹⁰ In Table 4, we test the overall significance of the regression for each country and each measure with the F-statistic. In Table 5, we report for each country and each measure the F-statistic and the p-value of the hypothesis that the monetary overhang measure can be excluded from the equation (the parameters of the lag polynomial $\beta_{\hat{\varepsilon}}(L)$ are all zero).

First, the results of overall significance tests of the forecasting models show that inflation is predictable at any horizon between one month and two years. As expected, it is easier to forecast inflation at long-run horizons (1 to 2 years) than in the short run (1 to 3 months). The results also seem to indicate that the performances of the two models are very similar on short horizons. However, some differences appear for long horizons.

⁹ It is well known, ever since Friedman and Schwartz (1963), that the delays in monetary policy transmission to output and then to prices are very long, situated between one to two years. Belongia and Ireland (2016) confirm the existence of such delays and suggest that they could be even longer today. *A priori*, we can think that it is more easy to predict inflation over a period of one or two years than on a horizon of one or three months. Carstensen et al. (2009) also take into account in their application, a forecast horizon of up to two years.

¹⁰ Carstensen et al. (2009) set the lag order to four, because their data are quarterly.

Table 4. Tests of the overall significance of the inflation forecasting model (F-statistic)

	Hungary	Czech Republic	Poland
h = 1			
MO based on OEM	5.229 (0.000)	6.849 (0.000)	3.571 (0.000)
MO based on CEM	5.260 (0.000)	6.750 (0.000)	3.416 (0.000)
h = 3			
MO based on OEM	10.164 (0.000)	11.475 (0.000)	7.962 (0.000)
MO based on CEM	10.011 (0.000)	11.477 (0.000)	7.614 (0.000)
h = 12			
MO based on OEM	19.845 (0.000)	13.339 (0.000)	9.408 (0.000)
MO based on CEM	17.566 (0.000)	13.642 (0.000)	8.779 (0.000)
h = 24			
MO based on OEM	22.281 (0.000)	16.257 (0.000)	7.537 (0.000)
MO based on CEM	18.463 (0.000)	17.511 (0.000)	6.779 (0.000)

Note: P-values of the hypothesis of nullity of the parameters of $\beta_{\pi}(L)$ and $\beta_{\varepsilon}(L)$ are reported in brackets.

Table 5. Tests of exclusion of the monetary overhang (MO) measures in the forecasting regressions

	Hungary	Czech Republic	Poland
h = 1			
MO based on OEM	1.195 (0.285)	0.723 (0.738)	0.611 (0.842)
MO based on CEM	1.234 (0.258)	0.620 (0.835)	0.399 (0.968)
h = 3			
MO based on OEM	1.818 (0.043)	0.261 (0.995)	1.085 (0.374)
MO based on CEM	1.681 (0.068)	0.263 (0.995)	0.744 (0.717)
h = 12			
MO based on OEM	3.183 (0.000)	0.898 (0.556)	1.590 (0.092)
MO based on CEM	1.816 (0.044)	1.105 (0.358)	1.026 (0.428)
h = 24			
MO based on OEM	4.027 (0.000)	1.409 (0.160)	2.077 (0.018)
MO based on CEM	1.908 (0.033)	2.146 (0.014)	1.302 (0.217)

Note: P-values of the hypothesis of nullity of the parameters of $\beta_{\varepsilon}(L)$ are reported in brackets.

Next, the results of the monetary overhang exclusion tests are globally consistent with the results of the cointegration tests. For the Czech Republic, monetary overhang measures are not good predictors of future inflation, whatever is the forecasting horizon (except for the monetary overhang based on the CEM, which predicts inflation at a horizon of two years). This result is

not surprising, as we show that the hypothesis of stationarity of monetary overhang is doubtful in the case of the Czech Republic. For Hungary and Poland, the monetary overhang based on OEM gives a better prediction of future inflation at all horizons than the monetary overhang based on CEM (in the sense that the probabilities are lower). Inflation seems to be unpredictable (by the monetary overhang) at a very short horizon (one month). At a horizon of one quarter, monetary overhang shows a significant ability to predict inflation in the case of Hungary. Further, at a horizon of one and two years, we cannot exclude for Hungary and Poland the monetary overhang from the forecasting regression, especially when the OEM-based measure is used. The monetary overhang measure using CEM is a significant predictor of inflation in the case of Hungary, but not in the case of Poland.

To gain further insights in the forecasting ability of monetary overhang measures, we follow Stock and Watson (1999) and Carstensen et al. (2009), and perform forecasting comparisons using a recursive out-of-sample methodology for the last five years of the sample (from 2012-01 to 2016-11). For each model, the prediction error at horizon h , given a forecast carried out at date t , is:

$$e_{t,h} = \pi_{t+h}^h - f_{t,h}, \quad (05)$$

where $f_{t,h}$ is the forecast based on equation (04). The root mean-squared prediction errors (RMSPE) at horizon h are defined as:

$$RMSPE(h) = \sqrt{\frac{\sum_{t=1}^N (e_{t,h})^2}{N}} \quad (06)$$

Table 6 displays the RMSPE for the two models, for each country, and for $h = 1, 3, 12$ and 24. First, we notice for all countries that the RMSPE is a decreasing function of h up to a horizon $h = 12$, and then it decreases for Hungary and Poland when h increases from 12 to 24. Whatever the model, the best predictions of inflation are obtained with a horizon of a year for Czech Republic and Poland. For Hungary, the best predictions are obtained with a horizon of one or two years, depending on the model. Finally, when we consider the out-of-sample comparison of CEM-based and OEM-based forecasting regressions, the RMSPE indicates that, for Hungary and Poland, the best predictor is based on OEM (respectively at a horizon of two years and one year), whereas for Czech Republic it is based on the CEM (at a horizon of one year).

Table 6. Root mean-squared prediction errors (RMSPE)

	Hungary	Czech Republic	Poland
h = 1			
MO based on OEM	0.0374	0.0302	0.0206
MO based on CEM	0.0383	0.0301	0.0204
h = 3			
MO based on OEM	0.0262	0.0178	0.0163
MO based on CEM	0.0275	0.0182	0.0166
h = 12			
MO based on OEM	0.0179	0.0146	0.0112
MO based on CEM	0.0200	0.0131	0.0120
h = 24			
MO based on OEM	0.0174	0.0169	0.0143
MO based on CEM	0.0204	0.0139	0.0162

These results are consistent with those of the cointegration tests (Table 3): if the hypothesis of cointegration is doubtful (i.e., for the Czech Republic), the RMSPE indicates that the OEM-based money overhang does not improve the forecast of inflation vis-à-vis the CEM-based monetary overhang. However, comparison of point estimators (RMSPE) is only indicative, and we cannot conclude that the forecast accuracy of the two models is significantly different. In addition, we cannot be sure that one measure does not contain information that is already contained in the other.

To test the hypothesis of equal forecast accuracy, we follow Harvey et al. (1997) and we use the MDM statistic, which is an adjustment for moderate-sized sample of the DM statistic of Diebold and Mariano (1995).

We denote d_t as the loss differential $d_t = (e_t^{CEM})^2 - (e_t^{OEM})^2$ where e^{CEM} and e^{OEM} are the two forecast errors associated with the two models. Diebold and Mariano (1995) propose a test for the null hypothesis of equal forecast accuracy based on the sample mean loss differential \bar{d} . Their test statistic is $DM = \bar{d}/\sigma_{\bar{d}}$ where $\sigma_{\bar{d}}$ is a consistent estimator of the standard deviation of \bar{d} , robust to heteroscedasticity and autocorrelation of order $h-1$. The asymptotic distribution of DM is the standard normal distribution. Harvey et al. (1997) propose the modified test statistic $MDM = N^{-1/2}[N + 1 - 2h + N^{-1}h(h - 1)]^{1/2}DM$, where N is the number of forecast, which corrects for small sample bias. For Harvey et al. (1997, 1998), it is very doubtful that the forecast errors are normally distributed and they recommend using the Student's t distribution with (N-1) degrees of freedom rather than the standard normal distribution to test the null hypothesis.

Table 7 presents the normality test of the forecast errors based on the Jarque-Bera statistics. Contrary to Harvey et al.'s (1997) assumption, the hypothesis of normality of the forecast errors is highly believable and there is no evidence that the forecast error distribution

is heavy-tailed (unless in the case of the Czech Republic at a horizon of one month). Therefore, we do not follow the recommendation of Harvey et al. (1997, 1998) to substitute the Student's t distribution to the normal distribution in testing for equal forecast accuracy (Table 8) or for forecast encompassing (Table 9).¹¹

Table 7. Normality test of the forecast errors (Jarque-Bera test)

	Hungary	Czech Republic	Poland
<hr/>			
h = 1			
OEM	0.062 (0.969)	12.258 (0.002)	1.966 (0.374)
CEM	0.016 (0.991)	13.861 (0.001)	1.050 (0.591)
<hr/>			
h = 3			
OEM	1.365 (0.505)	0.727 (0.695)	3.818 (0.148)
CEM	0.956 (0.619)	0.692 (0.707)	2.270 (0.321)
<hr/>			
h = 12			
OEM	4.222 (0.121)	0.611 (0.736)	3.696 (0.157)
CEM	3.995 (0.135)	0.533 (0.766)	3.469 (0.176)
<hr/>			
h = 24			
OEM	3.418 (0.180)	1.294 (0.523)	1.291 (0.524)
CEM	3.465 (0.176)	1.376 (0.502)	2.063 (0.356)

Note: The p-values of the normality tests are in bracket below the Jarque-Bera statistics.

Table 8 presents the results of the tests of equal forecast accuracy of the two forecasting models for each forecasting horizon, for the three countries. The (absolute) value of the MDM statistic is low and it is never above the critical value at the 10% level. For each country and each forecasting horizon, we cannot detect any notable difference in the forecast accuracy of the two models. However, this kind of test is characterized by a reduced power, which may prevent us from detecting a moderate difference in the forecast accuracy of the models (Harvey et al., 1997, 1998).

Table 8. Tests of equal forecast accuracy (MDM statistic)

	Hungary	Czech Republic	Poland
h = 1	1.223	-0.455	-0.306
h = 3	1.206	0.972	0.693
h = 12	0.457	-1.593	1.409
h = 24	1.277	-1.538	1.510

Note: The MDM statistic is calculated as described by Harvey et al. (1997); it is positive (negative) if the RMSPE from the CEM-based forecasting model is higher (lower) than the RMSPE from the OEM-based version. As the test is two-sided, the absolute value of the MDM statistic is compared to the critical value, which is 1,645 at the 10% level.

¹¹ In any case, N might be considered large enough in our sample (N = 59) to replace the Student's t distribution with the normal one, without affecting the results.

In a complementary exercise, we test for forecast encompassing. Following Stock and Watson (1999), in order to detect a significant improvement of the forecasting, we define a forecast combination regression:

$$\pi_{t+h}^h - \pi_t = \lambda f_{t,h}^{CEM} + (1 - \lambda)f_{t,h}^{OEM} + \eta_{t+h}, 0 \leq \lambda \leq 1, \quad (07)$$

where $f_{t,h}^{OEM}$ and $f_{t,h}^{CEM}$ are the two competing forecasts.

Equation (07) can alternatively be written with the forecast errors:

$$e_{t,h}^{OEM} = \lambda(e_{t,h}^{OEM} - e_{t,h}^{CEM}) + \eta_{t+h}. \quad (08)$$

Estimation of equation (08) allows us to test for forecasting encompassing. The objective is to assess whether the forecast $e_{t,h}^{CEM}$ contains useful forecasting information that is not contained in the alternative forecast $e_{t,h}^{OEM}$. The null hypothesis is $\lambda = 0$, and the alternative is $\lambda > 0$. When the null hypothesis is not rejected, the forecast $f_{t,h}^{OEM}$ is said to encompass the forecast $f_{t,h}^{CEM}$. In addition, the ‘reverse’ model $e_{t,h}^{CEM} = \lambda(e_{t,h}^{CEM} - e_{t,h}^{OEM}) + \eta_{t+h}$ has to be estimated too, with the null hypothesis $\lambda = 0$, to be tested against $\lambda > 0$.

We first test these hypotheses with a t-statistic that is robust to heteroscedasticity and autocorrelation of order h-1. We also implement the MDM framework test of Harvey et al. (1998) for forecast encompassing. It is based on the two-loss differential $d_{1t} = e_t^{OEM}(e_t^{OEM} - e_t^{CEM})$ and $d_{2t} = e_t^{CEM}(e_t^{CEM} - e_t^{OEM})$, used respectively for testing whether f_t^{OEM} encompasses f_t^{CEM} and whether f_t^{CEM} encompasses f_t^{OEM} . The MDM test of Harvey et al. (2008) is based on the statistics $MDM = N^{-1/2}[N + 1 - 2h + N^{-1}h(h - 1)]^{1/2}DM$, where $DM = \bar{d}/\sigma_{\bar{d}}$ is successively calculated on the series $d_{1,t}$ and $d_{2,t}$, with a consistent estimator $\sigma_{\bar{d}}$ of the standard deviation of \bar{d} , robust to heteroscedasticity and autocorrelation of order h-1.

Table 9 presents the results of the encompassing tests, for each forecasting horizon, for the three countries. These results are consistent with those of Table 6, and offer a more precise view of the ranking of the forecast models than the results of the tests of forecast accuracy equality (Table 8). Moreover, the results seem to be robust as the two tests deliver the same conclusion, even if the MDM statistic exhibits a reduced power (for the benefit of a more stable size).

For the Czech Republic, the OEM-based model encompasses the CEM-based version (while the reverse is not true) at a horizon of three months. However, at longer horizons (characterized by lower RMSPE), it is the CEM-based model that encompasses the OEM-based model. For Hungary, at horizons of one month, three months and two years, the OEM-based

model encompasses the CEM-based version. In addition, for Poland, we notice the same phenomenon at horizons of three months, one year and two years.

Table 9. Encompassing tests

	Hungary	Czech Republic	Poland
<hr/>			
h = 1			
f_t^{CEM} encomp. f_t^{OEM}			
λ	2.099	-0.025	0.244
t-statistic	1.751**	-0.021	0.299
MDM	1.547*	-0.021	0.306
f_t^{OEM} encomp. f_t^{CEM}			
λ	-1.099	1.025	0.755
t-statistic	-0.917	0.876	0.928
MDM	-0.869	0.896	0.853
<hr/>			
h = 3			
f_t^{CEM} encomp. f_t^{OEM}			
λ	1.983	1.445	0.936
t-statistic	2.130**	1.687**	1.560*
MDM	1.481*	1.397*	1.416*
f_t^{OEM} encomp. f_t^{CEM}			
λ	-0.983	-0.445	0.063
t-statistic	-1.056	-0.519	0.106
MDM	-0.871	-0.483	0.102
<hr/>			
h = 12			
f_t^{CEM} encomp. f_t^{OEM}			
λ	1.289	-0.725	1.291
t-statistic	1.120	-1.549	3.931***
MDM	0.669	-1.057	1.873**
f_t^{OEM} encomp. f_t^{CEM}			
λ	-0.289	1.725	-0.291
t-statistic	-0.251	3.684***	-0.887
MDM	-0.189	2.013**	-0.643
<hr/>			
h = 24			
f_t^{CEM} encomp. f_t^{OEM}			
λ	1.786	-1.837	1.878
t-statistic	2.865***	-3.079	13.079***
MDM	1.496*	-1.322	1.510*
f_t^{OEM} encomp. f_t^{CEM}			
λ	-0.786	2.837	-0.878
t-statistic	-1.261	4.756***	-6.116
MDM	-0.857	1.715**	-1.488
<hr/>			

Notes: (i) The t-statistic is used for testing the null hypothesis $\lambda = 0$ (the hypothesis of encompassing is not rejected) against the alternative hypothesis $\lambda > 0$ (the hypothesis of encompassing is rejected). (ii) The MDM statistic is calculated as described by Harvey et al. 1998). (iii) Both statistics are robust to heteroscedasticity and autocorrelation of order h-1. (iv) ***, **, * means significance at 1%, 5% and 10% significance level.

In short, for the CEE countries that exhibit a stable long-run money demand function, including a significant currency substitution effect (that is for Hungary and Poland), the OEM-based monetary overhang improves the forecast of inflation, comparatively to the CEM-based money overhang, in the sense that the first forecast encompasses the second. For the CEE

country that does not present a stable long-run money demand (the Czech Republic), the consideration of currency substitution does not improve the forecast of inflation.

5. Robustness analysis

Our findings may, however, be influenced by the time span we refer to. As shown by Figures 1-3, at the beginning of 2000's, the monetary overhang has a different path for the considered CEE countries. However, the trend is very similar after their EU accession. This evidence might influence the forecasting results of our models. Therefore, we conduct a similar investigation for the period 2004 to 2016, where we check whether the monetary overhang is a good predictor of inflation and whether there are significant differences between the OEM and CEM.

We first start with the test of the overall significance of the inflation forecast model (Table 10). The results show that the inflation forecasts improve with the increase in time horizon, up to one year for Hungary and Poland. Second, the tests for exclusion of the monetary overhang measures in the forecasting regressions (Table 11) are in line with previous findings, stating that the OEM-based monetary overhang is a good predictor of inflation for Hungary, but not for the Czech Republic. However, a slight difference appears in this case of Poland, where different from the main results, we notice that the monetary overhang is not a good predictor of inflation for any models and time-horizons under consideration.

Table 10. Tests of the overall significance of the inflation forecasting model (2004-2016)

	Hungary	Czech Republic	Poland
h = 1			
OEM	4.558 (0.000)	5.390 (0.000)	2.842 (0.000)
CEM	4.610 (0.000)	5.114 (0.000)	2.856 (0.000)
h = 3			
OEM	9.289 (0.000)	8.487 (0.000)	6.042 (0.000)
CEM	9.034 (0.000)	8.293 (0.000)	6.155 (0.000)
h = 12			
OEM	18.563 (0.000)	9.302 (0.000)	11.101 (0.000)
CEM	18.392 (0.000)	13.291 (0.000)	11.799 (0.000)
h = 24			
OEM	15.763 (0.000)	10.181 (0.000)	7.498 (0.000)
CEM	14.712 (0.000)	18.197 (0.000)	7.944 (0.000)

Note: P-values of the hypothesis of nullity of the parameters of $\beta_{\pi}(L)$ and $\beta_{\varepsilon}(L)$ are reported in brackets.

Table 11. Tests of exclusion of the monetary overhang in the forecasting regressions (2004-2016)

	Hungary	Czech Republic	Poland
h = 1			
MO based on OEM	1.278 (0.236)	1.097 (0.368)	0.394 (0.969)
MO based on CEM	1.337 (0.201)	0.818 (0.640)	0.411 (0.963)
h = 3			
MO based on OEM	2.520 (0.004)	0.581 (0.864)	0.516 (0.910)
MO based on CEM	2.311 (0.009)	0.441 (0.950)	0.616 (0.836)
h = 12			
MO based on OEM	3.470 (0.000)	0.371 (0.975)	0.864 (0.592)
MO based on CEM	3.833 (0.000)	2.881 (0.001)	1.273 (0.240)
h = 24			
MO based on OEM	2.569 (0.004)	0.278 (0.993)	0.586 (0.859)
MO based on CEM	2.044 (0.025)	4.574 (0.000)	0.896 (0.559)

Note: P-values of the hypothesis of nullity of the parameters of $\beta_{\xi}(L)$ are reported in brackets.

We continue the robustness check analysis and we perform forecasting comparisons using a recursive out-of-sample technique. The RMSPE is reported in Table 12. As in the previous case, the RMSPE is a decreasing function of h up to a horizon $h = 12$, for all three countries. In addition, the OEM-based forecasting regressions slightly outperform the CEM-based regression for Hungary and Poland, for at horizons of one month, three months and two years.

Table 12. Root mean-squared prediction errors (2004-2016)

	Hungary	Czech Republic	Poland
h = 1			
OEM	0.0354	0.0296	0.0179
CEM	0.0353	0.0294	0.0180
h = 3			
OEM	0.0244	0.0172	0.0144
CEM	0.0247	0.0174	0.0145
h = 12			
OEM	0.0173	0.0129	0.0113
CEM	0.0171	0.0117	0.0107
h = 24			
OEM	0.0201	0.0128	0.0105
CEM	0.0204	0.0098	0.0130

In what follows, we test for equal forecast accuracy, following Harvey et al. (1997). We first show that the standard normal distribution should be used to test the null hypothesis (Table 13).

Table 13. Normality test of the forecast errors (2014-2016)

	Hungary	Czech Republic	Poland
h = 1			
OEM	0.555 (0.757)	5.904 (0.052)	0.529 (0.767)
CEM	0.764 (0.682)	8.939 (0.011)	0.318 (0.852)
h = 3			
OEM	0.842 (0.656)	0.571 (0.751)	0.640 (0.726)
CEM	0.813 (0.665)	0.723 (0.696)	0.345 (0.841)
h = 12			
OEM	1.056 (0.589)	2.914 (0.232)	3.667 (0.159)
CEM	1.024 (0.599)	0.275 (0.871)	4.180 (0.123)
h = 24			
OEM	3.228 (0.199)	2.382 (0.303)	0.324 (0.850)
CEM	3.288 (0.193)	0.039 (0.980)	2.317 (0.313)

Note: The p-values of the normality tests are in bracket below the Jarque-Bera statistics.

Second, we present the results of the tests of equal forecast accuracy for the two forecasting models (the MDM statistic). Similar to the main results, we notice that the (absolute) value of the MDM statistic is lower than the critical value, showing no significant difference between the retained models (Table 14).

Table 14. Tests of equal forecast accuracy (2004-2016)

	Hungary	Czech Republic	Poland
h = 1	-0.216	-0.341	1.107
h = 3	0.825	0.279	0.700
h = 12	-0.215	-0.530	-0.453
h = 24	0.337	-0.455	-0.243

Note: Please refer to Table 8.

For the last step, we test for forecast encompassing, considering our reduced sample (Table 15). While the main results show that CEM-based model encompasses the OEM-based model for the Czech Republic, and the reverse applies for Hungary and Poland, the robustness check shows no significant difference between the CEM and OEM. These findings confirm the

equal forecast accuracy results (MDM statistics) but contrast the outcomes of recursive out-of-sample comparisons and tests for the exclusion of the monetary overhang.

Table 15. Encompassing tests (2004-2016)

	Hungary	Czech Republic	Poland
<hr/>			
h = 1			
f_t^{CEM} encomp. f_t^{OEM}			
λ	0.255	0.246	2.970
t-statistic	0.227	0.326	1.414
MDM	0.228	0.321	1.305
f_t^{OEM} encomp. f_t^{CEM}			
λ	0.744	0.753	-1.970
t-statistic	0.663	1.000	-0.938
MDM	0.647	1.034	-0.898
<hr/>			
h = 3			
f_t^{CEM} encomp. f_t^{OEM}			
λ	1.597	0.673	2.246
t-statistic	1.229	1.128	1.002
MDM	1.201	1.061	0.879
f_t^{OEM} encomp. f_t^{CEM}			
λ	-0.597	0.326	-1.246
t-statistic	-0.459	0.547	-0.555
MDM	-0.447	0.528	-0.511
<hr/>			
h = 12			
f_t^{CEM} encomp. f_t^{OEM}			
λ	-0.263	0.231	-0.574
t-statistic	-0.091	0.502	-0.275
MDM	-0.074	0.444	-0.233
f_t^{OEM} encomp. f_t^{CEM}			
λ	1.263	0.768	1.574
t-statistic	0.438	1.668	0.755
MDM	0.357	1.048	0.688
<hr/>			
h = 24			
f_t^{CEM} encomp. f_t^{OEM}			
λ	1.973	0.295	-0.200
t-statistic	0.752	1.421	-0.110
MDM	0.449	1.196	-0.067
f_t^{OEM} encomp. f_t^{CEM}			
λ	-0.973	0.704	1.200
t-statistic	-0.371	3.392	0.662
MDM	-0.223	0.936	0.392
<hr/>			

Notes: Please refer to Table 9.

To sum up, the OEM-based monetary overhang improves the inflation forecasts for Hungary and Poland, and, in particular, at long horizons. This result is not confirmed for the Czech Republic, where the stability of money demand function is put into question. In addition, the discrepancies between the two models for estimating the stability of money demand decrease after their EU accession. Our results partially contrast the findings reported by Horvath et al. (2011), who state that money matters but does not improve the predictability of inflation

in the CEE countries. However, our findings explain why studies that use closed-economy models for investigating the money demand in CEE countries (e.g. Narayan, 2010) report more cases of unstable money demand, while researchers who rely on open-economy specification, show that the money demand in CEE countries is, in general, stable (e.g. Dreger et al., 2007; Fidrmuc, 2009).

6. Conclusions and policy implications

A stable money demand function shows to what extent the monetary aggregates can be used in the conduct of monetary policy, and it allows us to test whether the monetary overhang is a good predictor of future inflation. Moreover, a stable money demand shows that a monetary-targeted regime is feasible to meet the Maastricht criteria for the CEE countries.

To provide additional insights to the money demand stability in selected CEE countries, we compare two money demand functions, resulting from a closed and an open, micro-founded money demand model. This specification allows us to see whether the consideration of a currency substitution effect makes the money demand more stable in the long run. Using monthly data for the Czech Republic, Hungary and Poland for the period 1999-2016, our results can be summarized as follow.

First, considering the interest rate differential (OEM) we obtain consistent results for money demand stability, although to a smaller extent for the Czech Republic. This result might be explained by the fact that the structure of money in circulation in the Czech Republic (approximated through the structure of bank deposits), shows a higher proportion of domestic currency denominated deposits as compared to Hungary and Poland (around 90% of deposits in the Czech Republic compared to 80% in the other two countries). The policy implications of these findings show that in countries where the confidence in domestic currency is higher, the effect of currency substitution for monetary stabilization diminishes.

Second, we discover that the monetary overhang generated from the OEM surpasses the monetary overhang that is estimated from the CEM, especially in the pre-crisis period. On the one hand, this observation points in favor of a soft monetary stance before the crisis. On the other hand, this evidence raises questions about the forecasting capacity of the monetary overhang. Given the structural break that appears in the monetary overhang in 2008 for all the countries retained in the analysis, we admit that our forecast exercises have some limits.

However, our F-test confirms the cointegration results and states that the monetary overhang is a good predictor of inflation only for Hungary and Poland.

Third, the results of the forecasting exercises are mixed and we cannot detect a significant difference of forecast accuracy between the two competing models. On the one hand, we have performed out-of-sample inflation forecasts. These results show that prediction errors decrease with the time horizon, until 12 months for all three countries, results confirmed by our robustness analysis. Nevertheless, the consideration of the monetary overhang in forecasting the inflation is recommend only for Poland, and to a smaller extent for Hungary. Poland, as the other countries in our sample, has in place an inflation-targeting monetary strategy and a floating exchange rate regime. However, as compared to the Czech Republic and Hungary that have a trade openness of over 160% of GDP according to the World Bank statistics in 2014, the trade openness in Poland is about 90% of its GDP. This evidence partially contradicts the important role of international factors in explaining the money demand in Poland. However, the central bank of Poland announced an accommodating inflation targeting policy in order to sustain economic growth and to achieve financial stability. Therefore, money becomes more important in predicting inflation if monetary policy becomes less restrictive, but helps to predict inflation only in the long run.

On the other hand we have tested the equal forecast accuracy of OEM- and CEM-based models. The MDM statistics (Harvey et al., 1997) does not highlight any notable difference in the forecast accuracy of the two models. This result also is confirmed by the robustness check we have performed for the period 2004-2016. Finally, we have tested for the forecast encompassing using a forecast combination regression. In this case, the main findings show that for the Czech Republic, the CEM-based model encompasses the OEM-based model at long horizons. An opposite situation is recorded for Hungary and Poland at horizons of three months, one year and two years. However, the robustness check for encompassing tests shows no significant difference between the two models.

All in all, we can conclude that: (i) the open-economy specification of the money demand model gives more consistent results than the closed-economy version for Hungary and Poland; (ii) the ability of the OEM-based monetary overhang to better predict inflation can be met only for Hungary and Poland, and increases with the time horizon (for these countries money matters in predicting inflation); and (iii) we cannot detect a significant difference of forecast accuracy between the two competing models, especially when we consider the post-EU accession period.

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Appendixes

Appendix A - Data description

Variables	Database	Explanations
m = M/P (real money)	OECD IFS (IMF)	M - Broad Money (M3) Index (2010=100), seasonally adjusted. P - Consumer Prices Index (2010=100).
y (output)	IFS (IMF)	Industrial Production Index (2010=100), seasonally adjusted.
r (domestic interest rate)	Eurostat	1-month Money Market Rate for the Czech Republic, Hungary and Poland.
r* (Euro Area interest rate)	Eurostat	1-month Money Market Rate for the Euro Area.

Appendix B – Stability of money demand (robustness analysis)

Table B1

Estimation results (2004-2016)

	Hungary		Czech Republic		Poland	
	OEM	CEM	OEM	CEM	OEM	CEM
α_0	-2.040* (-1.854)	-1.703* (-1.706)	-7.977*** (-14.377)	-6.710*** (-8.815)	-6.410*** (-19.395)	-6.808*** (-16.016)
α_1	0.436* (1.914)	0.371* (1.787)	1.712*** (14.552)	1.455*** (8.959)	1.412*** (21.029)	1.497*** (17.419)
α_2	-0.011 (-1.478)	-0.011 (-1.620)	-0.008 (-0.809)	-0.057*** (-4.304)	-0.038*** (-4.705)	-0.023** (-2.400)
α_3			0.148*** (7.548)		0.016** (2.372)	

Notes: (i) The t-statistics are in bracket below the estimated long run parameters; (ii) ***, **, * means significance at 1%, 5% and 10% significance level.

Table B2

Tests of the hypothesis $\alpha_1 = 1$ (2004-2016)

	Hungary		Czech Republic		Poland	
	OEM	CEM	OEM	CEM	OEM	CEM
t-statistic (probability)	-2.475 (0.014)	-3.030 (0.002)	6.054 (0.000)	2.804 (0.005)	6.139 (0.000)	5.785 (0.000)

Table B3

Results of the cointegration tests (Hansen's parameter instability tests) (2004-2016)

	Hungary		Czech Republic		Poland	
	OEM	CEM	OEM	CEM	OEM	CEM
Lc statistic (probability)	1.161 (< 0.01)	1.113 (< 0.01)	0.846 (0.046)	0.421 (0.072)	0.514 (> 0.2)	0.364 (0.106)