The Economic Research Guardian – Vol. 5(1)2015
Semi-annual Online Journal, www.ecrg.ro
ISSN: 2247-8531, ISSN-L: 2247-8531
Econ Res Guard 5(1): 73-96

ON THE LONG RUN MONEY–PRICES RELATIONSHIP IN CEE COUNTRIES

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Abstract
We investigate the role of money in explaining the long run inflation in 12 CEE countries, using monthly data for the period 2004-2013. We use a panel cointegration approach and recently developed empirical techniques as the panel fully modified and the panel dynamic regression procedures. Beside the role of interest rate and economic growth rate in explaining inflation, our cointegration equation explores the role of broad money (M2) growth. We also look to the M2 components, namely M1 and the difference between M2 and M1. We find no cointegration relationship either for the broad money or for the narrow money. However, money created by the banking sector explains the inflation in CEE countries in the long run. This last finding characterizes the entire panel, the panel of the seven CEE countries candidates to the Euro area, but not the panel of Euro area members. The findings are robust regarding the consideration of the income velocity's impact on the money in circulation.

Keywords: Money, inflation, Panel cointegration, CEE countries

JEL classification: E52, E31, E51, E40, C23

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1. Introduction

While the idea that inflation is associated with the growth of money is one of the oldest propositions in macroeconomics (for a historical perspective on this relationship, see Păun and Topan, 2013), the role of money in monetary policy conduct has been largely disputed during the last decades. The unconventional monetary policies implemented by central banks all over the world after the setup of the recent financial crisis, raised once again the interest for the investigation of the money–prices relationship. This issue is even more stringent for the Central and Eastern European (CEE) countries exposed to the European Central Bank (ECB) monetary policy, and where monetary easing measures were consistently applied even before the crisis. Consequently, a natural question arises: does the excess liquidity endanger price stability in transition economies in the long run?

Macroeconomists have constantly observed that prolonged increases in prices are associated with increases in the nominal quantity of money (Zhang, 2013). Starting with Friedman (1963), this link has been documented several times, inter alia by Brillembourg and Khan (1979), Lucas (1980), Gerlach (2004), Aksoy and Piskorski (2006), Nelson (2008) and Sargent and Surico (2008). However, before the crisis, the mainstream viewed little interest in analyzing money developments (see, e.g. Woodford, 2003, 2008). In this line De Grauwe and Polan (2005), Stavrev (2006) or Roffia and Zaghini (2007) suggest that the importance of money for inflation may be limited.

In order to restart the credit activity and the economy, central banks have increased the monetary base several times since the crisis outbreak (Croitoru, 2013). For this purpose a series of non-standard, unconventional monetary policy (UMP) measures were taken by the main central banks, including quantitative easing measures, as lending to financial institutions, liquidity injections and large-scale asset purchase programs (Roache and Rousset, 2013; Cerna, 2014). Nevertheless, these measures have not produced their effects on the economic growth. Moreover, the monetary policy rates are closed to zero. Against this background, two concerns raise the interest of monetary specialist. The first one refers to the relation between money and inflation while the second concern is related to the monetary policy effectiveness when the policy rate is equal to zero.

This paper addresses the first concern, with an application for CEE countries. The UMP measures undertaken by the national banks were considered necessary because the effectiveness of monetary policy was seriously reduced in most CEE countries in the context of the global financial crisis (ECB, 2010). In addition, the CEE countries members of the Euro area have benefit directly from the liquidity created by the ECB. Furthermore, several CEE countries (i.e. Hungary and Poland) established agreements on repurchase transactions with the ECB (IMF, 2013). Consequently, two situations may appear in relation with the quantitative easing in CEE countries. On the one hand, if the additional money pass-through inflation, the economic recovery depends of the firms' response to the quantitative easing and a risk of hard-to-control inflation levels appears. On the other hand, if the relationship between money and inflation cannot be documented in the long run, then a liquidity trap situation exists.

There is enough room for studying the link between money and prices in CEE countries. Two recent papers addressing this issue investigate the money demand function (Fidrmuc, 2009), and the role of money in forecasting inflation (Horváth et al., 2011). Both papers employ cointegration techniques and panel regression techniques that take into account the potential endogeneity of the involved variables, as the fully modified OLS estimator (FMOLS) and the dynamic OLS (DOLS) estimator.

The case of CEE countries needs to be further explored for several reasons. First, several countries encountered high inflation rates during the 1990s when they had in place monetary aggregate target regimes, or fixed exchange rate regimes, so the quantity of money in circulation meters. Starting with the second half of the 1990s, countries as the Czech Republic, Poland, Hungary, Slovak Republic and Romania passed to inflation targeting regimes (Minea and Tapsoba, 2014). Given the effects of the recent global financial crisis, there are several challenges to central bank monetary policies like increased uncertainty, unconventional measures or the link between inflation targeting and financial stability (Horváth et al., 2011). Second, few countries joined the Euro area, while other countries shall join the Exchange Rate Mechanism II (ERM II). Therefore, a stable money demand and an effective monetary policy transmission mechanism represent necessary pre-conditions for the introduction of euro (Elbourne and de Haan, 2006).

We will essentially pursuing a similar objective as Fidrmuc (2009) and Horváth et al., (2011), but differently from these previous papers which focus on the determination of long run money demand stability, we regress inflation on money growth (for a discussion on the interest of the two approaches regarding quantity-theory relations between money growth and inflation, see McCallum and Nelson, 2010). In this line, money growth represents an information variable of inflation, beside the real interest rate and the economic growth.

The money–prices relationship is one of the central focuses of the empirical literature on applied cointegration analysis (Budina et al., 2006). Cointegration implies that in the long run, different nonstationary variables should not move too far away from each other (Choudhry, 1995). Consequently we test for the long run relationship between the growth rate of prices, the growth rate of money, the economic growth rate and the real interest rate, in a panel framework, including 12 CEE countries[1], out of which five are Euro area members. We test this relationship for the period 2004-2013 and we adopt a cointegration analysis for heterogeneous panels, as proposed by Pedroni (1999, 2001). Even if the CEE countries can be seen as a homogenous group, strong heterogeneity persist between these countries, part of them being Euro area members, while the candidates ones have implemented different monetary policy regimes. Consequently, the adoption of cointegration techniques conceived for homogenous panels as in (Kao, 1999) can conduct to biased results (see Hurlin and Mignon, 2007).

Another contribution of our paper to the exiting literature consists in the use of monthly data, while previous studies analyzing money–prices relationship employ quarterly data, considering the economic growth restriction. In practice the monetary policy decisions are taken on a monthly basis. Thus, the use of monthly data seems appropriate and for doing so, previous researches (i.e. Fidrmuc,

[1] Bulgaria, Cyprus, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Romania, Slovak Republic, Slovenia.

2009) employ the industrial production indicator as a proxy for the economic growth. However, this indicator presents a high variability and considerably influences the empirical findings. As an alternative, we transform the quarterly economic growth data in monthly data using a cubic spline function[2] and afterwards we seasonally adjust this series.

Furthermore, we investigate the money–prices relationship with a reference to the broad money (M2 monetary aggregate) but we also look to the M2 components, namely M1 and the non-M1 deposits (that is, the difference between M2 and M1). This investigation is important on the one hand due to the fact that M1 and M2 series may move differently (McCallum and Nelson, 2010) and, on the other hand, it is stringent to see if money created by central banks explains the long run inflation, or the money created by commercial banks have this characteristic.[3]

Finally, our analysis makes abstraction of the money velocity in the first step, similar to cash-in-advance models. This abstraction is commonly met in researches which investigate the long run co-movement between money and prices as described by the quantity theory of money.[4] However, in reality the velocity fluctuates as shown by Baumol (1952). In the same vain, the Friedmanite proposal on the link between the income velocity of money and money growth, states that the velocity influences the variability of money growth. Consequently, in the second step we correct the money growth in respect of the influence of the velocity. The stationary test performed on the income velocity of money computed based on the quantity theory documents the non-stationarity of this indicator. Consequently, making abstraction of the velocity's influence on the money growth, can lead to inconsistent results.

The rest of the paper is structured as follows. Section 2 presents a brief review of the literature on money–prices relationship. Section 3 describes the methodology while Section 4 presents the data and the results. Section 5 performs a robustness investigation of the money–prices relationship in CEE countries, considering the influence of money velocity. The final section concludes.

2. Literature review

The past decades have witnessed a real flood of econometric literature dealing with the money–prices relationship, where the cointegration techniques have represented the workhorse of the research.

Even if the interest for money in explaining the inflation was never lost, the mainstream mode of policy analysis before the crisis, frequently does not consider monetary aggregates in theoretical and

[2] The application of a quadratic function yields similar results.
[3] In the Post-Keynesian theory of money endogeneity, money is said to originate as bank-created loans from deposits, which in turn create more loans under the endogenous money supply (Badarudin et al., 2013).
[4] Estrella and Mishkin (1997) argue that in low inflation environments the velocity shocks are usually accommodated by central banks and have no implication for inflation.

empirical investigations (see, i.e. Woodford, 2000; Svensson, 2008). As Thornton (2014) shows, "money's role in monetary policy has been tertiary, at best". The mainstream considers that central banks affect economic activity and inflation through the short term interest rate and by influencing the economic agents' expectations. However, the interest for exploring the money role in explaining the prices has increased once again in the aftermath of the financial crisis, given the plethora of quantitative easing measures adopted by the major central banks.

The exploration of money–prices relationship almost necessarily begins with a discussion of the quantity theory of money (QTM). Consequently, a first line of research is dedicated to the validation or invalidation of the equation of exchange (Friedman's quantity theory of money). The actual form of this identity shows that if a change in the quantity of (nominal) money were exogenously engineered by the monetary authority, then the long run effect would be a change in the price level (McCallum and Nelson, 2010).

Recent works challenging this theory are those of Zhang (2013) and Graff (2013). The first study investigates the monetary dynamics of inflation in China using a multivariate cointegration analysis and the vector error-correction (VEC) system developed by Johansen (1991, 1995). Its results suggest that inflation in China is Granger-caused by monetary growth in both the short and the long run. The second study challenges the QTM for a set of 109 countries from 1991 up to 2012 using a panel analysis and a fixed-effects model. The author notes that the classical proportionality theorem does not hold but excess money growth is a reliable predictor of inflation.

A second line of research on the money–prices relationship addresses the money demand function, which is one of the more heavily investigated subjects in macroeconomics. This equation is mainly tested in the case of emerging economies. In this line, Choudhry (1995) shows the stationarity of the long run money demand function in Argentina, Israel, and Mexico, using the Johansen technique. In a similar vein, Price and Nasim (1999) find a well-defined money demand relationship for Pakistan. More recently, Pelipas (2006), using cointegration analysis and dynamic equilibrium correction models, discovers that a long run function for real money balances exists for Belarus.

The investigation of the long run money–prices relationship frequently makes abstraction of the income velocity of money, which is assumed to be constant. However, changes in money velocity exist since movements of money supply do not always correspond to a matching money demand. Consequently, another strand of literature focuses on modeling money velocity. Rodríguez Mendizábal (2006) assess the behavior of money velocity in high and low inflation countries, based on a general equilibrium model of money demand where the velocity of money is determined endogenously as in Baumol (1952). A rather different approach is adopted by Baunto et al. (2011) which test the Friedman's (1984) theory regarding the link between the decline in money velocity and the increased variability of the growth rate of money supply in Philippines. Finally, El-Shagi and Giesen (2013) estimate the "money overhang" in the US based on the multivariate state space model of velocity. They analyze the short run impact of money on prices and provide evidence for a substantial impact in the US.

The money–prices relationship is frequently used to explain the differences between low and high inflation regimes. In this line, Basco et al., (2009) focus on the Argentina's case and use a cointegration analysis. They show that proportionality holds for the high inflation period but weakens once inflation lowers. More recently, Amisano and Fagan (2013) employ a Markov Switching model for inflation with time varying transition probabilities for the Euro area, Germany, the US, the UK and Canada. They discover that that a smoothed measure of broad money growth has important leading indicator properties for switches between inflation regimes.

The analysis of the money–prices relationship in Europe is not intensively debated. The literature focuses on the information content of money in forecasting inflation. In this line, Trecroci and Vega (2000) and Nicoletti-Altimari (2001) show among others that monetary aggregates are the best forecasters of inflation in the Euro area. Berger and Österholm (2008) use in their turn a mean-adjusted Bayesian VAR model as an out-of-sample forecasting tool of Euro area inflation and document strong evidence that including money improves forecasting accuracy. On contrary, Horváth et al. (2011) analyzing the transition countries case, find that money matters for future inflation at the policy horizons that central banks typically focus on but money does not in general improve the inflation forecasts. Differently from the previous studies, Stavrev and Berger (2012) use a money-based New Keynesian Dynamic Stochastic General Equilibrium (DSGE) models and find that incorporating money in the model provides better performance than its cashless counterpart. More recently, Dreger and Wolters (2014) examine the money demand for Euro area, investigating in the same time the role of the broad monetary aggregate (M3) in predicting Euro area inflation.

The analyses of the money–prices relationship in CEE countries address the money demand function and usually employ cointegration technique. On the one hand, there are studies which investigate a single-country case. In this vain, Budina et al. (2006) explore the Cagan money demand function for Romania. Similar, Vizek and Broz (2009) examine the effect of excess money growth on inflation in Croatia. Finally, Păun and Topan (2013) analyze the relationship between broad money dynamics and the CPI inflation using a simple VAR technique, in order to illustrate the monetary causes of inflation in Romania.

On the other hand, a series of papers address the case of CEE countries in a panel framework. For example, using panel cointegration techniques and different estimators, Fidrmuc (2009) find that money demand in selected CEE countries is significantly determined by the euro area interest rates and the exchange rate against the euro, which indicates a possible instability of this function.

We proceed to a similar analysis without modeling the money demand function in CEE countries. We consider the money growth, as well as the real interest rate and the economic growth rate as determinants of inflation (for an explanation, see Hall et al., 2009). We adopt a cointegration technique specific for heterogeneous panels and we provide evidence on the causal relationship between prices, money, interest rate and economic growth in 12 CEE countries, using Dumitrescu and Hurlin (2012) Granger non-causality test for heterogenous panels.

3. Methodology

3.1. Panel Cointegration Tests

View the heterogeneity of the CEE countries retained in our panel, we assume the presence of dynamic heterogeneity across groups. Consequently, our analysis relies on the Pedroni (1999, 2001) panel cointegration tests. Different from the Kao (1999) test constructed for strictly homogenous panels, the Pedroni's tests allows for cross-section interdependence with different individual effects and relax the homogeneity assumption. The general equation is:

$$Y_{i,t} = \propto_{i,t} + \delta_i t + \gamma_{1,i} X_{1i,t} + \gamma_{2,i} X_{2i,t} + \cdots + \gamma_{2,i} X_{2i,t} + \varepsilon_{i,t} \quad (1)$$

where $i = 1, \ldots, N$ for each country in the panel; $t = 1, \ldots T$ refers to the time period; parameters $\propto_{it}, \delta_i t$ allow the existence of country-specific fixed effects and deterministic trends; $\varepsilon_{i,t}$ denote the residuals, associated with deviations from the long run relationship.

The null hypothesis of no cointegration ($\rho_i = 1$) is tested applying a unit root test on the residuals, as follows:

$$\varepsilon_{i,t} = \rho_i \varepsilon_{it-1} + w_{it} \quad (2)$$

Pedroni (1999, 2001) proposes two sets of tests for cointegration, one for homogenous panels and the other for heterogeneous ones. The first category of tests is based on the within dimension approach (panel statistics) which includes four statistics: panel v-Statistic, panel rho-Statistic, panel PP-Statistic, and panel ADF-Statistic. These statistics pool the autoregressive coefficients across different countries for the unit root tests on the estimated residual. The group tests are based on the between dimension approach which includes three statistics: group rho-Statistic, group PP-Statistic and group ADF-Statistic. The between-group estimator is less restrictive because nonparametric tests have particular strengths when the data have significant outliers (for a description, see Hurlin and Mignon, 2007).

For the estimation of the cointegration relationship in our study, we propose a modified version of Pedroni's cointegration model (eq. 1), as follows:

$$cpi_{i,t} = \propto_{i,t} + \delta_i t + \gamma_{1,i} mmrr + \gamma_{2,i} gdp + \gamma_{3,i} M + \varepsilon_{i,t} \quad (3)$$

where cpi is the CPI inflation rate, $mmrr$ is the money market real rate, gdp is the economic growth rate and the M is the money growth rate.

3.2. Estimating the long run cointegration relationship by FMOLS and DOLS

After the documentation of the cointegration relationship, we first estimate the FMOLS technique for heterogeneous cointegrated panels, following Pedroni (2000). The model is:

$$cpi_{i,t} = \propto_{i,t} + \gamma_{1,i} mmrr_{i,t} + \gamma_{2,i} gdp_{i,t} + \gamma_{3,i} M_{i,t} + \varepsilon_{i,t} \quad (4)$$

where cpi and $mmrr$ are cointegrated with the slope $\gamma_{1,i}$, cpi and gdp are cointegrated with the slope $\gamma_{2,i}$, and cpi and M with the slope $\gamma_{3,i}$.

The FMOLS estimator is a non-parametric estimator. Alternatively, building up on Kao and Chiang (2000), Pedroni also propose a between-dimension, group means panel DOLS estimator that incorporates corrections for endogeneity and serial correlation parametrically. This is done by modifying (eq.1):

$$cpi_{i,t} = \propto_{i,t} + \gamma_{1,i} mmrr_{i,t} + \sum_{k=-k_i}^{k_i} \beta_{1ik} \Delta mmrr_{it} + \gamma_{2,i} gdp_{i,t} + \sum_{k=-k_i}^{k_i} \beta_{2ik} \Delta gdp_{it} + \gamma_{3,i} M_{i,t} + \sum_{k=-k_i}^{k_i} \beta_{3ik} \Delta M_{it} + \varepsilon_{i,t} \quad (5)$$

Both approaches take into account the potential endogeneity of variables. On the one hand, Pedroni (2000)'s FMOLS corrects for the endogeneity and serial correlation to the OLS estimator non-parametrically. On the other hand, the DOLS estimator uses the future and past values of the differenced explanatory variables as additional regressors (Fidrmuc, 2009). Although the accuracy of the two models is comparable, we employ both models for robustness purpose.

3.3. Panel Granger Causality tests

After the identification of the cointegration relationship between variables, it is usefully to see the causal relationship between them. For our estimations we expect to have bi-directional causality relationship between all variables. For this purpose, we resort to a classic Pairwise Granger causality test for panel data, as well as the Dumitrescu and Hurlin (2012) Panel Causality Test for heterogenous panels.

The investigation of the Granger causality for the long run relationship is based on a two-step process. The first step is the estimation of the residuals from the long run model while the second is the estimation of the predicted residuals as a dependent variable in a dynamic error correction model.

4. Data and results

4.1. Data

For the 12 CEE countries, monthly data regarding CPI inflation, money and money market rate are collected from the IMF (International Financial Statistics) and cover the period 2004M12–2013M11.[5] The money growth rate is obtained relative to the same month of the previous year. The real money market rate is obtained as the difference between the nominal rate and the inflation rate. The GDP growth rates are extracted from the Eurostat database, with quarterly frequencies. The quarterly data are transformed into monthly data using a cubic spline function with the last observation matched to the source data and afterwards seasonally adjusted using the additive method of X12-ARIMA procedure (the trend of the variables for the analyzed period is presented in Appendix A) .

In order to assess the panel cointegration, all series must have the same order of integration. Therefore, a series of well-known panel unit root and stationarity tests, from the first generation, are performed in order to check the integration order. These tests are adapted for heterogenous panels and their results are presented in Table 1.

Table 1 - First generation panel unit root and stationarity tests

Variables		Levin, Lin and Chu t*	Im, Pesaran and Shin W-stat	ADF–Fisher Chi-square	PP–Fisher Chi-square	Hadri Z-stat
Level	*cpi*	-1.177	-3.298***	45.62***	29.25	4.976***
	mmrr	-0.344	-2.937***	41.79**	42.27**	3.474***
	gdp	-3.427***	-4.347***	62.32***	24.62	4.372***
	M2	-1.312*	-0.376	18.40	12.62	13.91***
	M1	-1.443*	-2.038**	32.75	25.53	5.953***
	M2M1	-0.364	-0.280	23.39	15.51	12.11***
First difference	Δ*cpi*	-1.379*	-8.187***	119.3***	534.3***	-1.671
	Δ*mmrr*	-3.107***	-11.41***	182.6***	647.6***	-1.945
	Δ*gdp*	-3.305***	-4.741***	64.04***	51.23***	-1.497
	Δ*M2*	-2.874***	-9.125***	140.8***	741.9***	-0.865
	Δ*M1*	-1.569*	-9.370***	143.4***	784.7***	-0.677
	Δ*M2M1*	0.902	-8.663***	132.0***	754.6***	-0.554

*Notes: (1) *, ** and *** indicates the rejection of the null hypothesis at 1% , 5% and 10% significance level; (2) Levin, Lin & Chu t* assumes common unit root process; (3) Im, Pesaran and Shin W-stat, ADF–Fisher Chi-square and PP–Fisher Chi-square assume individual unit root process; (4) The null hypothesis of Hadri Z-stat is the panel stationarity; (5) denotes the first difference; (6) cpi is the inflation rate, mmrr is the money market real rate, gdp is the economic growth rate, M2 is the broad money growth rate, M1 is the narrow money growth rate and M2M1 is the growth rate of the difference between broad and narrow money.*

While the LLC test (Levin et al., 2002) and the IPS test (Im et al., 2003) provide mixed evidence regarding the presence of a unit root in level for the selected variables, the other tests, in particular

[5] Money market rate for Hungary is obtained from Eurostat. Day-to-day rates are considered as 1-month or 3-months series are characterized by a considerable number of missing data.

the Hadri (2000) Lagrange multiplier (LM) test document the non-stationarity of the variables in level. In first difference, all the variables are stationary, being I(1).

However, these classic panel unit root tests rely on the cross-sectional independence hypothesis, which is a very strong assumption, and therefore easily over-reject the null hypothesis of unit root in the presence of cross-sectional dependence. Thus, we can accept the results of panel unit root tests, only if the cross-sectional independence hypothesis is verified. In this line, we use a series of cross-sectional dependence tests (Friedman, 1937; Frees, 1995; Pesaran, 2004) for three different panels (considering M2, M1 or the difference between M2 and M1 – M2M1) and we present the results in Table 2. We notice that in all the cases the null of cross-sectional independence is rejected, which question the results of the first generation of panel unit root tests.

Table 2 - Cross-sectional dependence tests

Tests	M2	M1	M2M1
Friedman Chi-square	667.8 (0.00)	907.4 (0.00)	409.9 (0.00)
Frees Normal	3.236 (0.00)	5.732 (0.00)	1.684 (0.00)
Pearson CD Normal	44.78 (0.00)	58.50 (0.00)	21.56 (0.00)

*Notes: (1) *, ** and *** indicates the rejection of the null hypothesis at 1% , 5% and 10% significance level; (2) The null hypothesis for each tests is the cross-sectional independence; (3) Test statistic are reported and p-values in brackets; (4) A normal distribution was used to approximate the Frees' Q distribution.*

We therefore use the second generation Pesaran cross-sectional Augmented Dickey–Fuller (CADF) test, to check for the presence of panel unit roots (Table 3). Pesaran (2007) advanced a modified statistics based on the IPS test (Im et al., 2003), considering the average of the individual CADF tests. While in level most of variables present unit roots, in first difference all variables are I(1). We can proceed thus with the cointegration analysis.

Table 3 - Second generation panel unit root tests

CADF test	cpi	mmrr	gdp	M2	M1	M2M1
Level	-2.546	-2.298	-1.470	-1.986	-2.096	-1.971
	(0.00)	(0.02)	(0.89)	(0.22)	(0.11)	(0.24)
First difference	-5.597	-5.756	-5.882	-5.194	-5.312	-5.041
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)

Notes: (1) The null hypothesis for both tests is the presence of panel unit root; (2) p-values in brackets; (3) The cross-sectional ADF (CADF) test is proposed by Pesaran (2007) assuming cross-sectional dependence; (4) t-bar is reported and the p-values are in brackets; (5) 2 lags are used for the CADF test.

4.2. Cointegration results

We present three categories of results. In the first category, as most of previous studies, we consider the growth rate of the broad money (M2) in our cointegration equation. Afterwards, we decompose

M2 in its components, M1 and M2M1 (the difference between M2 and M1) and we compute the growth rate of these components.[6] Thus, we are able to see, on the one hand, if the choice of a specific monetary aggregate influences the results and, on the other hand, if the money created by the central banks (M1 in particular), or rather the money created by the banking sector (M2M1) influence the inflation in the long run.

The tests of Pedroni (2001) have seven different statistics for checking possibility of cointegration. The first four statistics are the panel cointegration statistics and are based on the within approach or homogenous cointegration. The other three tests, called group panel cointegration statistics, check heterogeneous cointegration. Additionally, there are four weighted statistics in the within dimension. A general cointegration relationship exists if the large part of tests documents this relationship. However, as we assume that our panels are heterogeneous due to the reasons explained above, we are interested in particular on the between dimension tests.

The first group of tests (Table 4) shows the three categories of results for all the selected CEE countries. For the M2, the homogenous tests of Pedroni (2001) indicate a long run relationship between inflation on the one hand, and the money market real rate, the GDP growth rate and the money growth rate. The results are not however confirmed when using weighted statistics. Moreover, the cointegration tests for heterogeneous panels do not show the existence of cointegration. We thus conclude that there is no long run relationship between our variables for the 12 CEE countries.

Table 4 - Pedroni (2001)'s panel cointegration tests (CEE-12)

	M2		M1		M2M1	
Within-dimension						
	Statistic	Weighted Statistic	Statistic	Weighted Statistic	Statistic	Weighted Statistic
Panel v-Statistic	1.753**	-0.057	0.206	-1.379	5.291***	1.686**
Panel rho-Statistic	-5.358***	0.391	-2.270***	2.376	-6.677***	-1.439*
Panel PP-Statistic	-5.955***	-0.241	-2.583**	2.582	-6.250***	-1.868**
Panel ADF-Statistic	-7.399***	-0.844	5.026	4.055	-7.602***	-1.311*
Between-dimension						
Between-dimension						
	Statistic		Statistic		Statistic	
Group rho-Statistic	0.064		2.241		-2.214**	
Group PP-Statistic	-0.486		3.069		-2.789***	
Group ADF-Statistic	-0.639		4.929		-2.516***	

*Notes: (1) *, ** and *** mean statistic relationship significant at 10%, 5%, 1%, respectively; (2) 1296 observations; (3) Akaike information criterion for lags selection is used.*

The absence of the cointegration relationship is evident when considering the growth rate of M1. So, the money created by central banks has no long run effect on inflation and thus, the quantitative easing measures undertaken recently do not affect the prices in the long run. The results are totally

[6] M3 data with monthly frequency are not available in the IMF statistics for CEE countries. In this case, M2 represents in our analysis the broad money.

different when we consider the long term deposits (M2M1). Their transformation in credits by the banking sector shows the long run relationship with the inflation, economic growth and real interest rate.[7]

In the second step we split our panel in two parts. The first part includes 7 CEE countries which are not yet Euro area member, while the second part contains the countries which have adopted euro. For the first group of countries, the results are reported in Table 5 and are similar with those obtained for the entire panel. In the case of M2 the cointegration relationship is uncertain, although the tests for homogenous panels confirm the long run co-movement of variables. However, when the M1 growth rate is retained into the analysis no cointegration relationship appears, while an opposite situation is documented for the difference M2M1.

Table 5 - Pedroni (2001)'s panel cointegration tests (CEE-7)

	M2		M1		M2M1	
Within-dimension						
	Statistic	Weighted Statistic	Statistic	Weighted Statistic	Statistic	Weighted Statistic
Panel v-Statistic	2.057**	-0.115	0.930	-0.970	5.814***	2.591***
Panel rho-Statistic	-6.117***	-0.152	-3.574***	1.822	-7.301***	-2.283**
Panel PP-Statistic	-6.112***	-0.798	-3.477***	1.895	-6.264***	-2.406***
Panel ADF-Statistic	-7.178***	-1.118	3.676	3.328	-7.061***	-1.716**
Between-dimension						
	Statistic		Statistic		Statistic	
Group rho-Statistic	-1.002		1.312		-3.264***	
Group PP-Statistic	-1.410*		2.162		-3.425***	
Group ADF-Statistic	-1.171		4.088		-3.063***	

*Notes: (1) *, ** and *** mean statistic relationship significant at 10%, 5%, 1%, respectively; (2) 756 observations; (3) Akaike information criterion for lags selection is used.*

The tests for the 5 CEE countries which are Euro area members (Table 6) show no cointegration between variables. The situation is the same, either if we consider the M2 or the M1 aggregate (or their difference), proving no long run link between the inflation rate and its determinants. The results are not surprising because the access to money created by the ECB can be limited to the small CEE countries. Furthermore, in the case of monetary aggregates and money market rate, we have the same observation for all the countries in the panel, after their accession to the monetary union. This situation can influence the results. All in all, we conclude that there is no cointegration relationship for the CEE-5.

[7] Our results might be affected by the presence of outliers in the sample. When we analyze the figures in Appendix A, we notice a strong variation in the growth rate of M1 for Romania, around the moment of its EU accession. Therefore, Romania can act as an outlier and influences the results. Consequently, for robustness purpose, we have eliminated Romania from the sample and we have performed once again the tests, for 11 CEE countries (Appendix B). The results are very similar with those reported in Table 4.

Table 6 - Pedroni (2001)'s panel cointegration tests (CEE-5)

	M2		M1		M2M1	
Within-dimension						
	Statistic	Weighted Statistic	Statistic	Weighted Statistic	Statistic	Weighted Statistic
Panel v-Statistic	-0.503	0.051	-1.265	-0.981	0.203	0.173
Panel rho-Statistic	1.152	0.817	1.868	1.526	-0.320	-0.098
Panel PP-Statistic	1.092	0.672	2.218	1.745	-0.601	-0.438
Panel ADF-Statistic	0.054	0.059	2.467	2.286	-1.549*	-0.279
Between-dimension						
	Statistic		Statistic		Statistic	
Group rho-Statistic	1.286		1.919		0.430	
Group PP-Statistic	0.914		2.195		-0.268	
Group ADF-Statistic	0.395		2.798		-0.273	

*Notes: (1) *, ** and *** mean statistic relationship significant at 10%, 5%, 1%, respectively; (2) 540 observations; (3) Akaike information criterion for lags selection is used.*

4.3. Estimating the long-run cointegration relationship by FMOLS by DOLS

After documenting cointegration relationships in the case of M2M1 for CEE-12 and CEE-7, we proceed to the computation of the individual estimators (Table 7). We notice that all variables have a strong impact of the inflation level and the sign is the expected one. First, the interest rate negatively influences the prices level, while the economic growth has an opposite impact. The increase in money created by the commercial banks positively influences the prices level in the long run. This evidence is stronger for CEE-7. Both FMOLS and DOLS estimators provide robust results.

Table 7 - Panel FMOLS and DOLS (M2M1)

variables	CEE-12		CEE-7		CEE-5	
	FMOLS	DOLS	FMOLS	DOLS	FMOLS	DOLS
MMRR	-0.803***	-0.825***	-0.854***	-0.883***	-	-
GDP	0.088***	0.089***	0.067***	0.062***	-	-
M2M1	0.095***	0.092***	0.101***	0.097***	-	-
R^2	0.740	0.794	0.737	0.807	-	-

*Notes: (1) *, ** and *** mean statistic relationship significant at 10%, 5%, 1%, respectively; (2) heterogeneous; (3) Akaike information criterion for lag and lead selection in the case of DOLS is employed.*

4.4. Panel Granger Causality tests

For the causality analysis, 2 lags have been chosen based on vector autoregressive (VAR) best lag order selection criteria. There are two categories of results, specific for homogenous panels and

The Economic Research Guardian – Vol. 5(1)2015
Semi-annual Online Journal, www.ecrg.ro
ISSN: 2247-8531, ISSN-L: 2247-8531
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considering common coefficients and for heterogeneous panels, considering individual coefficients (Table 8). The tests are performed for the CEE-12 group and for the CEE-7 group respectively.

Table 8 - Panel Causality Tests (M2M1)

M2M1	Granger causality test (common coefficient)			Dumitrescu Hurlin Panel Causality Tests (individual coefficient)		
	Null Hypothesis:	*F-stat.*	*(p-value)*	*Null Hypothesis:*	*Zbar-Stat.*	*(p-value)*
CEE-12	M2M1 does not Granger Cause CPI	9.205	(0.000)	M2M1 does not homogeneously cause CPI	4.186	(0.000)
	CPI does not Granger Cause M2M1	4.091	(0.016)	CPI does not homogeneously cause M2M1	6.084	(0.000)
	MMRR does not Granger Cause CPI	2.466	(0.085)	MMRR does not homogeneously cause CPI	3.818	(0.000)
	CPI does not Granger Cause MMRR	47.86	(0.000)	CPI does not homogeneously cause MMRR	5.538	(0.000)
	GDP does not Granger Cause CPI	31.54	(0.000)	GDP does not homogeneously cause CPI	9.882	(0.000)
	CPI does not Granger Cause GDP	3.923	(0.020)	CPI does not homogeneously cause GDP	3.309	(0.000)
	MMRR does not Granger Cause M2M1	0.339	(0.711)	MMRR does not homogeny. cause M2M1	3.805	(0.000)
	M2M1 does not Granger Cause MMRR	3.251	(0.039)	M2M1 does not homogeny. cause MMRR	2.817	(0.004)
	GDP does not Granger Cause M2M1	23.85	(0.000)	GDP does not homogeneously cause M2M1	13.77	(0.000)
	M2M1 does not Granger Cause GDP	0.477	(0.620)	M2M1 does not homogeneously cause GDP	-1.061	(0.288)
	GDP does not Granger Cause MMRR	7.164	(0.000)	GDP does not homogeneously cause MMRR	2.678	(0.007)
	MMRR does not Granger Cause GDP	0.334	(0.715)	MMRR does not homogeneously cause GDP	4.648	(0.000)
CEE-7	M2M1 does not Granger Cause CPI	10.75	(0.000)	M2M1 does not homogeneously cause CPI	3.938	(0.000)
	CPI does not Granger Cause M2M1	0.469	(0.625)	CPI does not homogeneously cause M2M1	3.116	(0.000)
	MMRR does not Granger Cause CPI	0.984	(0.374)	MMRR does not homogeneously cause CPI	1.246	(0.212)
	CPI does not Granger Cause MMRR	30.59	(0.000)	CPI does not homogeneously cause MMRR	5.562	(0.000)
	GDP does not Granger Cause CPI	19.23	(0.000)	GDP does not homogeneously cause CPI	7.537	(0.000)
	CPI does not Granger Cause GDP	1.992	(0.137)	CPI does not homogeneously cause GDP	2.911	(0.003)
	MMRR does not Granger Cause M2M1	0.585	(0.557)	MMRR does not homogeny. cause M2M1	1.500	(0.133)
	M2M1 does not Granger Cause MMRR	5.805	(0.003)	M2M1 does not homogeny. cause MMRR	2.831	(0.000)
	GDP does not Granger Cause M2M1	12.66	(0.000)	GDP does not homogeneously cause M2M1	9.370	(0.000)
	M2M1 does not Granger Cause GDP	0.030	(0.969)	M2M1 does not homogeneously cause GDP	-1.528	(0.126)
	GDP does not Granger Cause MMRR	5.148	(0.006)	GDP does not homogeneously cause MMRR	3.332	(0.000)
	MMRR does not Granger Cause GDP	0.097	(0.907)	MMRR does not homogeneously cause GDP	3.198	(0.001)

Notes: (1) 2 lags.

While the Pairwise Granger causality test shows mixed evidence, the test of Dumitrescu and Hurlin (2012), adapted for heterogeneous panels, documents the existence of a bi-directional causality relationship between variables, with few exceptions. For the first group of countries money Granger cause the level of inflation and a similar situation is obtained for the second group.

5. The influence of money velocity on the money growth rate

In Section 4 we have made abstraction about the role of income velocity of money, in influencing the money in circulation. However, the recent literature shows that money velocity influences the amount of money in circulation and thus the rate of inflation. For example, an increase in money supply implies a rise of uncertainty, which leads to a rise in money demand for precautionary reasons. Consequently, income velocity should decrease (see Baunto et al., 2011).

Even if most of papers addressing the long run relationship between money and prices consider the velocity as being constant, making abstraction of its impact without documenting its stationarity can

The Economic Research Guardian – Vol. 5(1)2015
Semi-annual Online Journal, www.ecrg.ro
ISSN: 2247-8531, ISSN-L: 2247-8531
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lead to biased results. Therefore, we have computed the income velocity of money based on the QTM and we have checked for the presence of unit roots (Table 9). Except for the LLC test, all other tests (from the first and second generation) show the presence of a unit root in the income velocity of money.

Table 9 - Panel unit root and stationarity tests for the velocity (in level)

Variables	Levin, Lin and Chu t*	Im, Pesaran and Shin W-stat	ADF–Fisher Chi-square	PP–Fisher Chi-square	Hadri Z-stat	CADF test
M2 velocity	-4.195***	-1.018	25.38	21.88	17.39***	-1.554
M1 velocity	-3.842***	-1.072	35.54*	26.52	14.69***	-1.509
M2M1 velocity	-1.260	0.945	12.81	10.69	10.81***	-0.999

*Notes: (1) *, ** and *** indicates the rejection of the null hypothesis at 1% , 5% and 10% significance level; (2) M1 velocity is computed by dividing the volume of the nominal GDP to the volume of M1, M2 velocity is computed by dividing the volume of the nominal GDP to the volume of M2 and the M2M1 velocity is computed by dividing the volume of the nominal GDP to the volume of the difference between M2 and M1; (3) For the Hadri test, the unit root is the panel stationarity; (4) 2 lags are used for the CADF test.*

This evidence forced us to correct the influence of velocity on the money in circulation. We have thus ortogonalized the money on the velocity using a simple OLS regression and we have retained the corrected form of the money growth. We have performed once again the tests for checking the robustness of the results presented in the previous section. Table 10 presents the cointegration results for all three panels: CEE-12, CEE-7 and CEE-5. For the CEE-12 group of countries and CEE-5, the results are similar. In the first panel, the cointegration relationship is documented only for the M2M1V. In the last panel, no cointegration relationship appears. So, we can conclude that our results are robust for these two panels.

However, in the case of the CEE-7 group, the cointegration relationship documented for the difference between M2 and M1 is not found anymore. The money velocity can thus interfere in estimation the long run money–prices relationship for these countries.

Table 10 - Pedroni (2001)'s panel cointegration tests (considering the influence of the money velocity)

	$M2_V$		$M1_V$		$M2M1_V$	
CEE-12						
Within-dimension						
	Statistic	Weighted Statistic	Statistic	Weighted Statistic	Statistic	Weighted Statistic
Panel v-Statistic	0.644	-0.849	0.457	-1.386	3.957***	0.908
Panel rho-Statistic	-4.236***	1.289	-2.395***	2.191	-5.382***	-0.938
Panel PP-Statistic	-4.733***	0.963	-2.906***	2.308	-5.389***	-1.439*
Panel ADF-Statistic	-5.088***	1.316	3.284	4.085	-4.707***	0.624
Between-dimension						
	Statistic		Statistic		Statistic	
Group rho-Statistic	0.914		1.711		-2.227**	
Group PP-Statistic	1.069		2.478		-2.685***	
Group ADF-Statistic	1.479		4.917		-0.961	
CEE-7						
Within-dimension						
	Statistic	Weighted Statistic	Statistic	Weighted Statistic	Statistic	Weighted Statistic
Panel v-Statistic	1.048	-0.753	1.241	-0.970	2.826***	2.354***
Panel rho-Statistic	-4.953***	1.051	-3.675***	1.893	0.224	0.263
Panel PP-Statistic	-4.968***	0.718	-3.757***	2.033	0.186	0.218
Panel ADF-Statistic	-5.605***	0.773	1.818	3.028	-0.854	-0.008
Between-dimension						
	Statistic		Statistic		Statistic	
Group rho-Statistic	0.155		1.105		1.184	
Group PP-Statistic	0.564		2.066		0.916	
Group ADF-Statistic	0.730		3.582		-0.395	
CEE-5						
Within-dimension						
	Statistic	Weighted Statistic	Statistic	Weighted Statistic	Statistic	Weighted Statistic
Panel v-Statistic	-0.828	-0.412	-1.295	-0.990	-0.578	0.039
Panel rho-Statistic	1.109	0.745	1.744	1.176	0.290	-0.411
Panel PP-Statistic	1.022	0.636	1.977	1.214	-0.137	-0.803
Panel ADF-Statistic	1.528	1.116	2.949	2.739	0.554	0.067
Between-dimension						
	Statistic		Statistic		Statistic	
Group rho-Statistic	1.233		1.343		-0.222	
Group PP-Statistic	0.989		1.395		-0.949	
Group ADF-Statistic	1.428		3.379		-0.073	

*Notes: (1) *, ** and *** mean statistic relationship significant at 10%, 5%, 1%, respectively; (2) Akaike information criterion for lags selection is used; (3) $M2_V$, $M1_V$ and $M2M1_V$ are the growth rate of monetary aggregates corrected for the influence of the money velocity.*

We continue with the individual estimators for the single cointegration relationship found in the case of the first panel. The signs are those expected and the influence of the interest rate, economic growth and money on prices is very strong. Furthermore, there is a good consistence between the results reported by the two estimators (Table 11).

Table 11 - Panel FMOLS and DOLS (considering the influence of the money velocity)

variables	CEE-12		CEE-7		CEE-5	
	FMOLS	DOLS	FMOLS	DOLS	FMOLS	DOLS
MMRR	-0.705***	-0.720***	-	-	-	-
GDP	0.118***	0.122***	-	-	-	-
$M2M1_V$	0.106***	0.106***	-	-	-	-
R^2	0.751	0.803	-	-	-	-

*Notes: (1) *, ** and *** mean statistic relationship significant at 10%, 5%, 1%, respectively; (2) heterogeneous; (3) Akaike information criterion for lag and lead selection in the case of DOLS is employed.*

Finally, we report the results of the Granger causality tests for panel data (Table 12). The first category of results shows the bi-directional causality in almost all the cases. The second set of results shows clearly that money cause inflation and vice-versa.

Table 12 - Panel Causality Tests (considering the influence of the money velocity)

$M2M1_V$	Granger causality test (common coefficient)			Dumitrescu Hurlin Panel Causality Tests (individual coefficient)		
	Null Hypothesis:	*F-stat.*	*(p-value)*	*Null Hypothesis:*	*Zbar-Stat.*	*(p-value)*
CEE-12	$M2M1_V$ does not Granger Cause CPI	3.433	(0.032)	$M2M1_V$ does not homogeneously cause CPI	3.839	(0.000)
	CPI does not Granger Cause $M2M1_V$	7.256	(0.000)	CPI does not homogeneously cause $M2M1_V$	6.828	(0.000)
	MMRR does not Granger Cause CPI	2.466	(0.085)	MMRR does not homogeneously cause CPI	3.818	(0.000)
	CPI does not Granger Cause MMRR	47.86	(0.000)	CPI does not homogeneously cause MMRR	5.538	(0.000)
	GDP does not Granger Cause CPI	31.54	(0.000)	GDP does not homogeneously cause CPI	9.882	(0.000)
	CPI does not Granger Cause GDP	3.923	(0.020)	CPI does not homogeneously cause GDP	3.309	(0.000)
	MMRR does not Granger Cause $M2M1_V$	0.234	(0.791)	MMRR does not homogeny. cause $M2M1_V$	1.673	(0.094)
	$M2M1_V$ does not Granger Cause MMRR	1.709	(0.181)	$M2M1_V$ does not homogeny. cause MMRR	3.923	(0.000)
	GDP does not Granger Cause $M2M1_V$	21.14	(0.000)	GDP does not homogeneously cause $M2M1_V$	12.01	(0.000)
	$M2M1_V$ does not Granger Cause GDP	1.778	(0.169)	$M2M1_V$ does not homogeneously cause GDP	-1.759	(0.078)
	GDP does not Granger Cause MMRR	7.164	(0.000)	GDP does not homogeneously cause MMRR	2.678	(0.007)
	MMRR does not Granger Cause GDP	0.334	(0.715)	MMRR does not homogeneously cause GDP	4.648	(0.000)

Notes: (1) 2 lags

Our analysis have however few limits. First, we have not considered the influence of the crisis. The identification of structural breaks in series for each of the country retained into analysis do not recommend the choice of a specific point in order to split the entire period in two sub-periods, in order to see if the crisis changed the long run money–prices relationship. Second, we have not considered the influence of different exchange rate regimes implemented in Euro area candidate countries. Third, we have assumed that our panels are heterogeneous but no tests were performed to prove the heterogeneity. Forth, in calculation of the inflation rate, no distinction between domestic and traded prices is made (see Price and Nasim, 1999).

6. Conclusions

The role of money in explaining the long run inflation in 12 CEE countries in investigated, in a panel cointegration framework. The money is considered as an inflation predictor, beside the real interest rate and the economic growth rate. The novelty of the paper consist in the decomposition of the broad money, which allow for testing the influence of money created by central banks and by the commercial banks respectively, on the inflation level.

Altogether, evidence whether a long run relationship exists between the selected variables, is documented only in the case of the difference between M2 and M1, for the CEE-12 and CEE-7 panels. We find no cointegration relationship for the broad money (M2), neither for the narrow money (M1), even if we correct for the influence of money velocity. However, when the cointegration relationship is found, the individual estimators and the panel Granger causality tests show a strong influence of money on prices in the long run.

The fact that money created by central banks has no long run influence on prices in the selected countries show the fact that the risk of hyperinflation which can be generated by the non-conventional monetary policy conducted recently by central banks does not exists. This is not surprising because, while money balances increased and nominal interest rates decreased in the period before the financial crisis, inflation did not accelerate at all (Dreger and Wolters, 2014).

References

Aksoy Y, Piskorski T (2006). U.S. domestic money, inflation and output. *Journal of Monetary Economics*. 53: 183-197.

Amisano G, Fagan G (2013). Money growth and inflation: A regime switching approach. *Journal of International Money and Finance*. 33: 118-145.

Badarudin ZE, Ariff M, Khalid AM (2013). Post-Keynesian money endogeneity evidence in G-7 economies. *Journal of International Money and Finance*. 33: 146-162.

Basco E, D'Amato L, Garegnani L (2009). Understanding the money–prices relationship under low and high inflation regimes: Argentina 1977–2006. *Journal of International Money and Finance*. 28: 1182-1203.

Baumol W (1952). The Transaction Demand for Cash: An Inventory-theoretic Approach. *Quarterly Journal of Economics*. 66: 545-556.

Baunto AL, Bordes C, Maveyraud S, Rous P (2011). Money growth and velocity with structural breaks: evidence from the Philippines. *Philippine Management Review*. 18: 71–81.

Berger H, Österholm P (2008) Does Money Growth Granger-Cause Inflation in the Euro Area? Evidence from Out-of-Sample Forecasts Using Bayesian VARs. *IMF Working Paper*, No. 53, International Monetary Fund.

Brillembourg A, Khan MS (1979). The Relationship between Money, Income and Prices: Has Money Mattered Historically. *Journal of Money, Credit and Banking*. 11: 358-365.

Budina N, Maliszewski W, De Menil G, Turlea G (2006). Money, inflation and output in Romania, 1992–2000. *Journal of International Money and Finance*. 25: 330-347.

Cerna S (2014). Unconventional Monetary Policy. *Revista OEconomica*. Issue 01.

Choudhry T (1995). High Inflation Rates and the Long-Run Money Demand Function: Evidence from Cointegration Tests. *Journal of Macroeconomics*. 17: 77-91.

Croitoru L (2013). What good is higher inflation? To avoid or escape the liquidity trap. Liquidity, the October 2008 speculative attack and central bank reputation. *NBR Occasional Paper*, No. 9, National Bank of Romania.

De Grauwe P, Polan M (2005). Is inflation always and everywhere a monetary phenomenon? *Scandinavian Journal of Economics*. 107: 239-259.

Dreger C, Wolters J (2014). Money demand and the role of monetary indicators in forecasting euro area inflation. *International Journal of Forecasting*. 30: 303-312.

Dumitrescu E-I, Hurlin C (2012). Testing for Granger non-causality in heterogeneous panels. *Economic Modelling*. 29: 1450-1460.

ECB (2010). The Impact of the Financial Crisis on the Central and Eastern European Countries. *ECB Monthly Bulletin*, July, European Central Bank.

Elbourne A, de Haan J (2006). Financial structure and monetary policy transmission in transition countries. *Journal of Comparative Economics*. 34: 1-23.

El-Shagi M, Giesen S (2013). Money and inflation: Consequences of the recent monetary policy. *Journal of Policy Modeling*. 35: 520-537.

Estrella A, Mishkin FS (1997). Is there a role for monetary aggregates for the conduct of monetary policy? *Journal of Monetary Economics*. 40: 279-304.

Fidrmuc J (2009). Money demand and disinflation in selected CEECs during the accession to the EU. *Applied Economics*. 41: 1259–-1267.

Frees EW (1995). Assessing cross-sectional correlation in panel data. *Journal of Econometrics*. 69: 393-414.

Friedman M (1937). The use of ranks to avoid the assumption of normality implicit in the analysis of variance. *Journal of the American Statistical Association*. 32: 675-701.

Friedman M (1963). *Inflation: Causes and Consequences*. Asia Publishing House.

Friedman M (1984). Lessons from the 1979-82 monetary policy experiment. *American Economic Review*. 74: 397-400.

Gerlach S (2004). The Two Pillars of the European Central Bank. *Economic Policy*. 19: 389-439.

Graff M (2013). *The Quantity Theory of Money in Year Six after the Subprime Mortgage Crisis*. In Maltritz D, Berlemann M (Eds.), Financial Crises, Sovereign Risk and the Role of Institutions, pp 169-195, Springer.

Hadri K (2000). Testing for stationarity in heterogeneous panel data. *Econometrics Journal*. 3: 148-161.

Hall SG, Hondroyiannis H, Swamy PAVB, Tavlas GS (2009). Assessing the causal relationship between euro-area money and prices in a time-varying environment. *Economic Modelling*. 26: 760-766.

Horváth R, Komárek L, Rozsypal F (2011). Does money help predict inflation? An empirical assessment for Central Europe. *Economic Systems*. 35: 523-536.

Hurlin C, Mignon V (2007). Une synthèse des tests de cointegration sur données de Panel. *Économie et Prévision*. 180: 241-265.

Im KS, Pesaran MH, Shin Y (2003). Testing for unit roots in heterogeneous panels. *Journal of Econometrics*. 115: 53-74.

IMF (2013). Global Impact and Challenges of Unconventional Monetary Policies. *IMF Policy Paper*, October, International Monetary Fund.

Johansen S (1991). Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models. *Econometrica*. 59: 1551-1580.

Johansen S (1995). *Likelihood-Based Inference in Cointegrated Vector Autoregressive Models*. Oxford University Press.

Kao C (1999). Spurious Regression and Residual-Based Tests for Cointegration in Panel Data. *Journal of Econometrics*. 90: 1-44.

Levin A, Lin CF, Chu CSJ (2002). Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties. *Journal of Econometrics*. 108: 1-24

Lucas R (1980). Two illustrations of the quantity theory of money. *American Economic Review*. 70: 1005-1014.

McCallum BT, Nelson E (2010). *Money and Inflation: Some Critical Issues*. In Friedman BM, Woodford M (Eds.), Handbook of Monetary Economics, Volume 3, pp. 97-153, Elsevier.

Minea A, Tapsoba R, (2014). Does inflation targeting improve fiscal discipline? *Journal of International Money and Finance*. 40: 185-203.

Nelson E (2008). Why money growth determines inflation in the long run: answering the Woodford critique. *Journal of Money, Credit and Banking*. 40: 1791-1814.

Nicoletti-Altimari S (2001). Does money lead inflation in the euro-area? *ECB Working Paper*, No. 63, European Central Bank.

Păun C, Topan V (2013). The monetary causes of inflation in Romania. *Romanian Journal of Economic Forecasting*. 16: 5-23.

Pedroni P (1999). Critical Values for Cointegration Tests in Heterogenous Panels with Multiple Regressors. *Oxford Bulletin of Economics and Statistics*. 61: 653-670.

Pedroni P (2000). Fully modified OLS for heterogeneous cointegrated panels. *Advanced in Econometrics*. 15: 93-130.

Pedroni P (2001). Purchasing power parity tests in cointegrated panels. *Revue of Economics and Statistics*. 83: 727-731.

Pelipas I (2006). Money demand and inflation in Belarus: Evidence from cointegrated VAR. *Research in International Business and Finance*. 20: 200-214.

Pesaran MH (2004). General Diagnostic Tests for Cross Section Dependence in Panels. *Cambridge Working Papers in Economics*, No. 0435, Faculty of Economics, University of Cambridge.

Pesaran MH (2007). A Simple Panel Unit Root Test in the Presence of Cross-Section Dependence. *Journal of Applied Econometrics*. 22: 265-312.

Price S, Nasim A (1999). Modelling inflation and the demand for money in Pakistan; cointegration and the causal structure. *Economics Modelling*. 16: 87-103.

Roache SK, Rousset MV (2013). Unconventional Monetary Policy and Asset Price Risk. *IMF Working Papers*, No. 190, International Monetary Fund.

Rodríguez Mendizábal H (2006). The Behavior of Money Velocity in High and Low Inflation Countries. *Journal of Money, Credit and Banking*. 38: 209-228.

Roffia B, Zaghini A (2007). Excess money growth and inflation dynamics. *ECB Working Paper*, No. 749, European Central Bank.

Sargent TJ, Surico P (2008). Monetary Policy and Low-Frequency Manifestations of the Quantity Theory. *External MPC Unit Discussion Paper*, No. 26, Bank of England.

Stavrev E (2006). Measures of underlying inflation in the euro area: assessment and role for informing monetary policy. *IMF Working Papers*, No. 197, International Monetary Fund.

Stavrev E, Berger H (2012). The information content of money in forecasting euro area inflation. *Applied Economics*. 44: 4055-4072.

Thornton DL (2014). Monetary policy: Why money matters (and interest rates don't). *Journal of Macroeconomics*. 40: 202-213.

Trecroci C, Vega JL (2000). The information content of M3 for future inflation. *ECB Working Paper*, No. 33, European Central Bank.

Vizek M, Broz T (2009). Modeling inflation in Croatia. *Emerging Markets Finance and Trade*. 45: 87-98.

Woodford M (2000). Monetary policy in a world without money. *International Finance*. 3: 229-260.

Woodford M (2003). *Interest and Prices: Foundations of a Theory of Monetary Policy*. Princeton University Press.

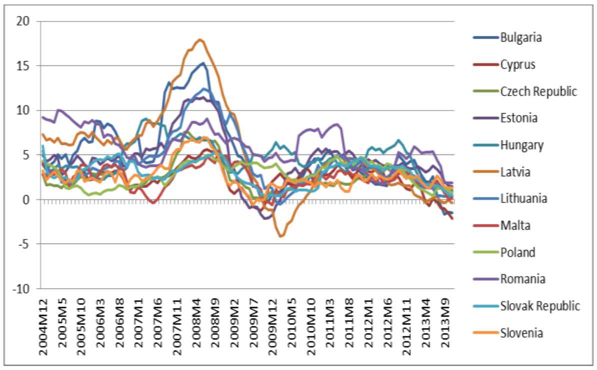
Woodford M (2008). How important is money for monetary policy conduct? *Journal of Money, Credit and Banking*. 40: 1561-1598.

Zhang C (2013). Monetary Dynamics of Inflation in China. *The World Economy*. 36: 737-760.

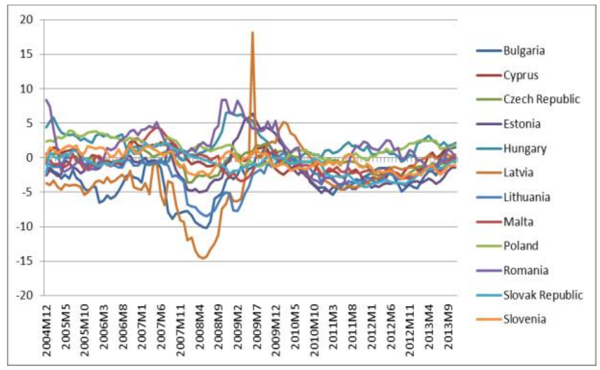
The Economic Research Guardian – Vol. 5(1)2015
Semi-annual Online Journal, www.ecrg.ro
ISSN: 2247-8531, ISSN-L: 2247-8531
Econ Res Guard 5(1): 73-96

Appendixes

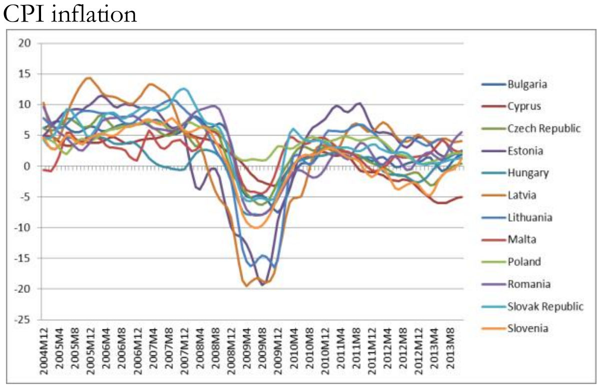
Appendix A - Stylized facts



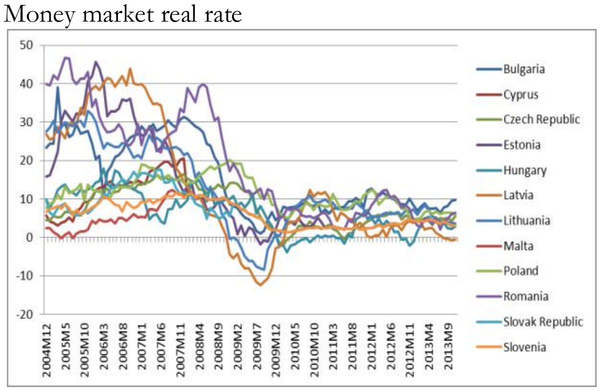
CPI inflation



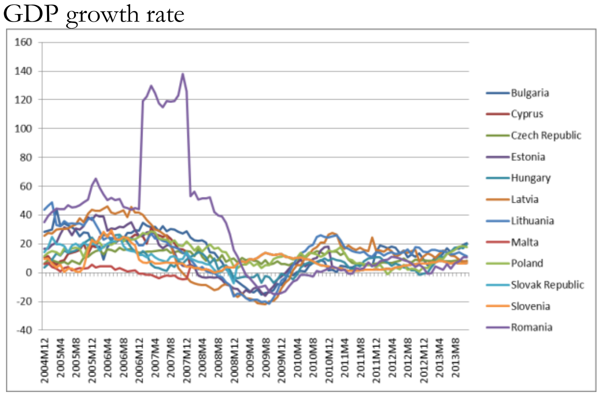
Money market real rate



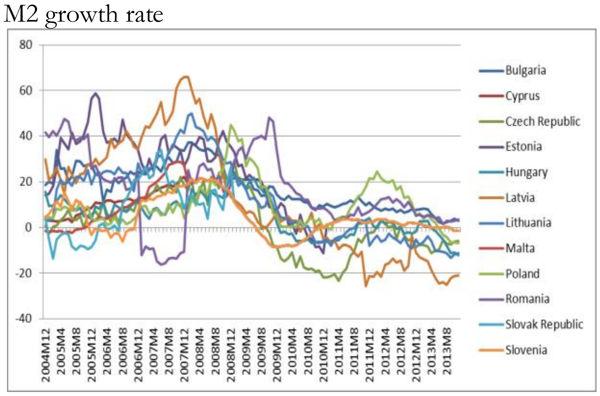
GDP growth rate



M2 growth rate



M1 growth rate



M2-M1 growth rate

Appendix B - Pedroni (1999)'s panel cointegration tests for CEE-11

	M2		M1		M2M1	
Within-dimension						
	Statistic	Weighted Statistic	Statistic	Weighted Statistic	Statistic	Weighted Statistic
Panel v-Statistic	1.952**	-0.091	0.221	-1.402	5.604***	1.534*
Panel rho-Statistic	-6.216***	0.404	-2.916***	2.288	-7.766***	-1.462*
Panel PP-Statistic	-6.435***	-0.131	-3.160***	2.467	-7.184***	-1.971**
Panel ADF-Statistic	-7.685***	-0.901	4.833	3.862	-8.440***	-1.587*
Between-dimension						
	Statistic		Statistic		Statistic	
Group rho-Statistic	-0.068		2.001		-2.420***	
Group PP-Statistic	-0.428		2.818		-3.069***	
Group ADF-Statistic	-0.925		4.557		-3.247***	

*Notes: (1) *, ** and *** mean statistic relationship significant at 10%, 5%, 1%, respectively; (2) 1188 observations*