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Contents

SUMMARY:* This paper undertakes an empirical investigation of the relationship between the dynamics of money supply and prices in Bulgaria over the period 1998–2012. Starting from the hypothesis that the endogeneity of money supply determines a link from prices (and possibly inflation) through money demand to money supply, the paper attempts to address the issue of the possibility of an existence of a feed-back effect of money aggregates (currency in circulation, M1 and M2) dynamics on inflation. The econometric analysis is based on three methods. The first one is dynamic cross-correlations which are performed after pre-whitening of the data and then applying univariate residual cross-correlation approach. The second one is Granger-causality tests in the framework of unrestricted VAR models which are conducted both on a rolling basis with a moving start date and fixed end date and with a moving window having a fixed length, with the aim of overcoming the problem of potential instability of coefficients. The third method is the Johansen cointegration technique which allows the analysis of both long-run relations between the variables, which are naturally related to the estimation of money demand models, as well as the short-run dynamics of money growth and inflation.

JEL Classification Numbers: E31, E41, C20, C30

Keywords: money supply, price level, inflation, univariate residual cross-correlation approach, Granger-causality tests, cointegrated VAR estimation, money demand models

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1. Introduction

This paper undertakes an empirical analysis of the relationship between the dynamics of money supply and prices in Bulgaria. The motivation of the paper derives from the fact that this topic has received relatively little empirical research since the period of the introduction of the currency board arrangement in the country, despite its importance for understanding money demand behaviour and the potentially interdependent relationship between money supply and prices/inflation. The analysis of money demand, that has been rather limited over the last years, is by itself important as it would allow a better understanding of the dynamics of monetary aggregates within the framework of the currency board and thus a deeper understanding of the monetary transmission mechanism in the country. In turn, an analysis of the dynamics of monetary indicators and prices could potentially extend the range of factors used for forecasting inflation, thus improving inflation forecasts. Taking a more general view, the relevance of the topic relates to the future membership of Bulgaria in the euro area where the ECB gives a prominent role to monetary aggregates under the money pillar of its monetary policy framework that is aimed at achieving the primary objective of maintaining price stability.

Money growth impacts inflation developments in the long run as advocated by the quantity theory of money, according to which money growth precedes equal changes in the general price level rate of growth. Nevertheless, the empirical evidence of the relationship between money growth and inflation and the usefulness of money for predicting inflation has been controversial, while the practical use of monetary indicators in the conduct of monetary policy by central banks across the world has also been subject to wide discussions and presently stands far from close to consensus.

The introduction of the currency board arrangement in Bulgaria in 1997 played a pivotal role in the macroeconomic stabilization of the country following the severe financial crisis that started in 1996. The strong nominal anchor of the fixed exchange rate significantly lowered inflationary expectations and eliminated one of the main contributors to inflation in the pre-currency board period – the large exchange rate fluctuations and their quick pass-through to the price level. Another important distinguishing feature of the currency board was the elimination of the possibility for providing credit to the government and the very limited potential opportunity for extending loans to commercial banks under the strict requirements of the lender of last resort function of the Bulgarian National Bank (BNB).

Theoretically, orthodox currency board arrangements bring about an automatic money supply adjustment mechanism. The latter constrains the growth of base money primarily to balance of payment dynamics and corresponding changes in money demand, thus making money supply endogenous. The currency board introduced in Bulgaria belongs to the so-called new generation (or non-orthodox) currency board arrangements which are characterized by the presence of some discretionary powers of monetary authorities that provide opportunities for deliberately affecting or intervening into money supply dynamics (see *e.g.* Nenovsky and Hristov, 2002). In the specific case of Bulgaria these discretionary powers relate to the minimum required reserves and the lender of last resort role of the BNB. The presence of certain channels of discretionary monetary policy especially the minimum required reserves implies that the expansion of money supply may deviate from that determined by market forces, *i.e.* from the automatic money supply mechanism.[1]

One of the aspects of the relationship between money supply and inflation in Bulgaria, related to the effects of past growth of money supply on current period inflation, has been investigated in a number of research papers in the years immediately following the introduction of the currency board, with most papers concluding there were no monetary sources of inflation. However, the short time period available at that time as well as the concentration of econometric methods primarily on single equations can be considered as limitations of these studies. This raises the need for a more thorough analysis of the possible link between money and prices since 1997, including the consideration of issues that have not been investigated deeply so far – both the short- and long-run relation between the variables as well as their potential two-way relationship with the application of multi-equation econometric methods such as VARs.

These specific research issues which lead to new conclusions from empirical econometric analysis for Bulgaria as well as the results from the cointegrated VARs related to the estimation of money demand models for currency in circulation, M1 and M2, which are published for the first time for Bulgaria, represent the original contributions of the paper. Furthermore, the empirical analysis in the paper spans a relatively long period of time:

[1] For an empirical investigation of the validity of the presence of an automatic mechanism under the currency board arrangement (*i.e.* reserve money dynamics following balance of payments dynamics), see Nenovsky, N. *et al.* (2001) for the case of Bulgaria, Estonia and Lithuania and Nenovsky, N. and K. Hristov (2002) only for the case of Bulgaria. Krus (2012) examines the effect of the current account on the changes in the monetary base for a sample of countries with currency board arrangements, using a panel regression method.

starting from 1998 (and depending on data availability from 2000) and ending in the first quarter of 2012.

The present paper analyses the following two hypotheses. The endogeneity of money supply mechanism and the driving force of money demand for money supply make us start with the hypothesis about a link that goes from prices (and inflation) through money demand to money supply. While the case of endogeneity might imply that money supply is an inappropriate factor in models explaining inflation developments, it could be also hypothesized that a feedback effect from money aggregates dynamics on inflation is also possible. For example, this feedback effect could be driven by factors that are external to the operation of the currency board itself (*e.g.* strong capital inflows may be regarded as possible monetary sources of inflation).

The second hypothesis is related to the question whether any short-run disequilibrium between money supply and demand that results in excess money supply could potentially act as a pro-inflationary factor in Bulgaria. This proposition is made originally by Sepp (1995). In his note on inflation under the Estonian currency board, Sepp maintains that while in the long-run money supply and money demand coincide, which practically excludes the possibility for money supply from being a long-run factor for inflation, over the short-run the situation may differ. In particular, in the short-run disequilibria between money supply and money demand may exist, with money supply potentially either constraining or fueling inflation (in the first case when money supply is temporarily lower than money demand and in the second case when money supply temporarily exceeds demand).

The empirical methodology that will be used to address the two research hypotheses – the potential two-way relationship between monetary aggregates and price developments and the effects from possible short-run disequilibrium between money supply and demand on inflation – is based on the following approaches. The initial statistical properties of the data will be investigated with dynamic cross-correlations, while the potential two-way causality will be analysed with the standard Granger-causality tests in a VAR framework. Empirical testing of the second hypothesis requires constructing money demand models which in itself represents an important contribution of the paper to the generally scarce literature on money demand, including an analysis on stability of money demand and especially over the period since 2000. To assess the validity of short-run excess of money supply having an effect on inflation and to allow treatment of potential endogeneity in the relationship between money and prices, we employ vector error-correction models (VECM). The latter represent an appropriate methodological approach as they make it possible to estimate a relationship between endogenous variables that is based both on their short-run dynamics and an adjustment to a possible long-run equilibrium.

The rest of the paper is structured as follows. In Section 2 we make a literature review of the research studies on money and prices/inflation, summarizing the results obtained so far for Bulgaria since 1998 as well as the studies for the countries with currency boards, specifically examining the Baltic States. Section 3 explores the time-series properties of the data on money and inflation, presents the methodological details of the applied econometric approach on dynamic cross-correlations and Granger-causality tests for the preliminary analysis of the data and reports the results. Section 4 presents the application of VECM in estimating money demand models for currency in circulation, M1 and M2. Section 5 summarizes the conclusions.

2. Literature Review

After the introduction of the currency board in Bulgaria a large number of studies were devoted to investigation of the impact of the new monetary policy framework on inflation. First, in 1998 Yotzov *et al.* report that over the first year of the operation of the currency board base money was not a determinant of inflation dynamics in contrast to the period before the currency board when the exchange rate, past inflation and base money represented the main inflation drivers. Yotzov (2000) finds that unlike the period from December 1991 to June 1997, in the period from July 1997 to December 1999 there was neither a short-term nor a long-term relationship between money supply (reserve money, M2 and M3) and inflation, based on the application of a single equation approach. Beck *et al.* (2003), who make an empirical assessment of the 'discipline' and 'confidence' effects of the established currency board on inflation, provide evidence that a confidence effect did occur, whereas the hypothesized lower effect of lagged money growth on the price level after the regime change found little support in inflation models. In a relatively recent study Arratibel *et al.* (2009) show that in an out-of-sample inflation forecasting exercise at a 12-quarter horizon over the period 2003Q3–2008Q2, single equation models for Bulgaria with monetary indicators included help outperform forecasts obtained with random walk models.

In the early period after the introduction of the new monetary policy regime there were a limited number of studies exploring money demand models for Bulgaria. Based on a single equation approach, Yotzov (2000) reports a strong statistical effect of real income on real money demand but does not succeed in identifying the theoretically expected negative association between interest rates and money demand, arguing that interest rates had not yet started functioning as an opportunity cost for holding money. In a research work exploring specifically transaction demand for currency in circulation, Nenovsky and Hristov (2000), also employing single equation modelling, conclude that demand for currency in circulation is most significantly determined by the monetary expenditures of households (retail sales in the economy being a weaker determinant). The authors also discover a statistically significant negative elasticity of demand with respect to the BGN/USD exchange rate, as an opportunity cost of holding currency, while the impact of interest rates on time deposits has the expected negative relationship with currency demand but is relatively low. Slavova (2003) estimated narrow and broad money demand models for Bulgaria for three periods: April 1991 – April 1996, May 1996 – June 1997, July 1997 – December 2000. The empirical analysis, conducted again within a single equation approach, reveals different determinants of money demand during the three periods. In the last period the author finds stabilization of money demand, with standard determinants of money demand such as the wage rate and the Treasury bill rate gaining statistically significant explanatory power.

The paper by Komárek and Melecký (2001) is one representative of a number of studies on money demand modeling, taking due consideration of factors specific for the economic development of transition countries. The authors construct a money demand model for narrow money in the Czech Republic over the period 1993–2001, using foreign variables such as the real effective exchange rate and foreign direct investment in addition to the traditional set of explanatory variables. In the case of broad money, the authors conclude that all standard domestic money demand factors apart from inflation have the expected theoretical impact on M2 and furthermore there is also a significant influence from currency substitution and capital mobility represented by the return on foreign deposits in USD and on US assets.

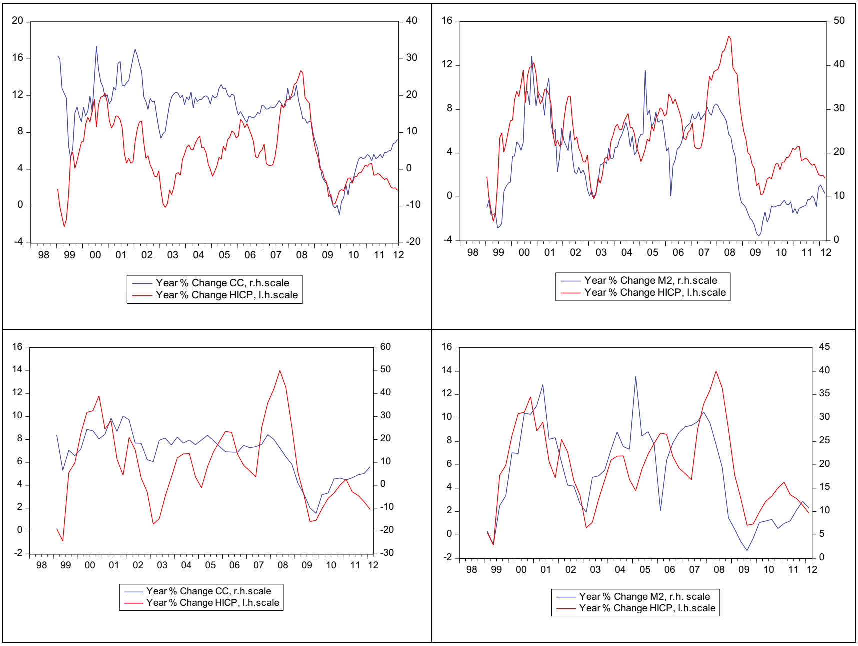
Pikkani (2000) specifies a model for the monetary sector in Estonia, modeling real M2 demand as a function of real GDP and a three-period moving average of inflation over the period from March 1995 to August 1999. Pikkani finds that income elasticity of money demand is higher than 1 (2.02). The explanation provided by the author is related to the potential role of money as an accumulator of wealth due to the practical absence of substitutes for money in the period of estimation. Another potential explanation is that income elasticity may also reflect initial conditions of low monetization, financial deepening, higher confidence and economic agents' optimistic expectations about the future. In a paper examining money demand in Latvia, Tillers (2004) establishes a stable money demand function using Johansen vector-error correction model for the period 1996–2003. The variables used for the analysis include real M2, real GDP and real long-term deposit interest rate. Siliverstovs (2007) constructs a money demand model for Estonia for the period 1995Q1–2006Q2 also with the Johansen cointegration procedure. He builds a model with real M2, real GDP, inflation rate as well as short- and long-term interest rates. The author discovers two cointegrating vectors: one characterizing a money demand function, and the other interpreted as a stationary spread between long- and short-term interest rates. As for the previous research for Estonia by Pikkani (2000) and the case of Latvia, the long-run income elasticity of money demand established by Siliverstovs is larger than unity (around 2).

The study by Dreger *et al.* (2006) investigates a long-run money demand relationship over the 1995–2004 period for 10 new EU Member States, using panel cointegration techniques. The authors do not manage to obtain such a relationship when employing the standard variables (broad money, real GDP and short-term market interest rate). Nevertheless, a stable long-run money demand cointegration link is established when the US dollar exchange rate is included. The authors find that income elasticity is larger than unity, while the elasticity with respect to the interest rate is negative with a relatively small magnitude.

Reimers and Roht (2007) construct a money demand function for Estonia using real M2 over the period from 1995 to 2006, applying various econometric techniques ranging from the system approach of Johansen to the Engle-Granger procedure and the dynamic OLS approach. The authors find unstable results when employing the Johansen approach but are able to obtain satisfactory results for a stable money demand function with the Engle-Granger approach. The preferred specification includes real GDP (with income coefficient greater than unity), euro area government bond rates with a maturity of 10 years and euro area money market rate.

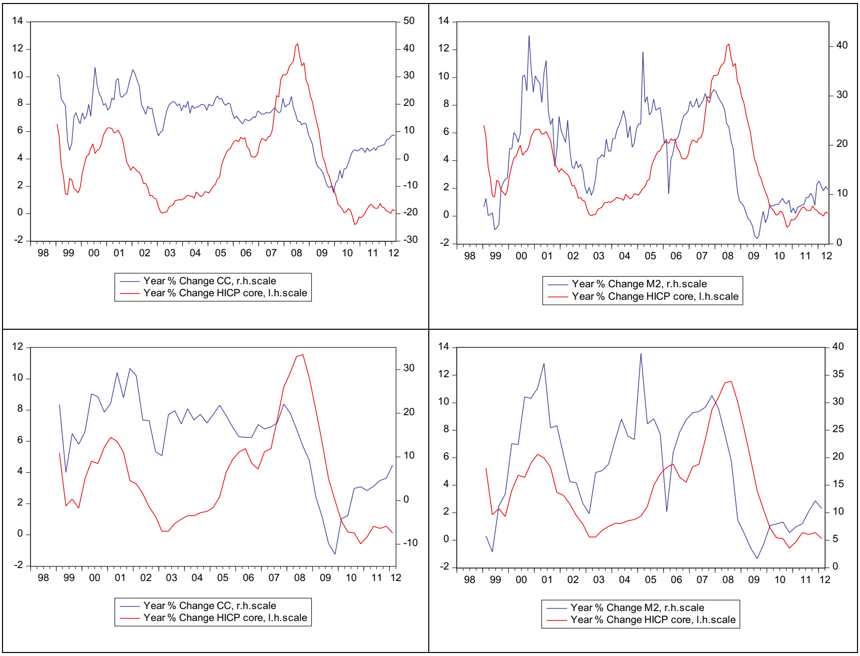
3. Analysis of the Statistical Properties of the Data

This section provides a preliminary analysis of the data focusing on their statistical properties over the period from January 1998 to March 2012, respectively for quarterly data from 1998Q1 to 2012Q1. We start with an investigation of the annual growth rates of monetary aggregates (currency in circulation, M1, M2) and the HICP index (overall and HICP excluding food, energy, administered prices and tobacco as a proxy for core inflation). Below we show the graphs for currency in circulation (CC) and M2 against overall HICP inflation. In the first line we employ monthly data, in the second line quarterly data respectively.



As seen from the graphs, despite the volatility of money growth and inflation, it appears that there is an association between the dynamics of the two series, perhaps more significantly for the case of M2 and HICP both when monthly and quarterly data are used. This finding seems to be confirmed for the developments of core inflation and the two selected monetary aggregates: currency in circulation and M2[2].

[2] The first line of the graphs presents monthly data and the second one quarterly data respectively.



Before proceeding with the empirical analysis, we seasonally adjust the following data series: currency in circulation (CC), monetary aggregate M1 (M1), monetary aggregate M2 (M2), the harmonized index of consumer prices (HICP) and HICP core (HICP_C). For the consumer price data series that are originally released on a monthly frequency, we transform them to quarterly by averaging, and for the monetary aggregates series we take an end-of-period transformation to obtain the respective quarterly frequency.

The preference for using quarterly rather than monthly data is determined by the following two reasons: 1) the substantial volatility present in the data with monthly frequency may be partly overcome with the quarterly aggregation. This argument makes it more appropriate to analyze dynamic cross-correlations of the data and Granger-causality tests based on data with quarterly frequency; 2) the empirical literature on the issue of cointegration reveals that cointegration depends on the total length of the sample rather than on the number of observations which does not give more power to statistical tests with monthly over quarterly data (see e.g. Otero and Smith (2000)).

The univariate time series properties of the seasonally adjusted data are examined by the Augmented Dickey-Fuller (ADF) test, the Phillips-

Perron (PP) test and the Kwiatkowski–Phillips–Schmidt–Shin (KPSS) test (see Appendix A1). The results point to the conclusion that the series are integrated of order one, even though for CC and M2 the tests are not conclusive. For robustness of the results, which are known to be characterized by low power especially in cases of structural breaks, we additionally employ the Clemente, Montañés and Reyes unit root test (1998) which can deal with two potential breaks in time series (see Appendix A2). The latter test reveals the presence of two structural breaks for most of the examined data (in some cases both in levels and in first difference form) and furthermore provides strong evidence for the treatment of all series as $I\,(1)$.

3.1. Dynamic Cross-correlations

The analysis of the relationship between monetary aggregates dynamics and inflation begins with dynamic cross-correlations that can help us to investigate the possible lead-lag links between the series.

The series are transformed by taking first differences of the respective log levels (on a quarterly basis, seasonally adjusted). When the series are transformed by differencing to ensure they are jointly covariance-stationary, then their interrelationship can be determined by examining either their cross correlation function or their cross spectrum. In this paper, we resort to the more frequently applied cross-correlation estimator. The cross-correlogram allows us to determine whether there is a one-way causality between the series or feedback occurring.

We use the following money aggregates: CC, M1 and M2, and the two price indicators: HICP and HICP_C. The table presented below shows the cross correlogram, *i.e.* the degree of co-movement of the first differences of the respective money and price indicators, that is estimated from 1998Q1 to 2012Q1. The correlation dynamics is estimated at lags and leads of 1, 2, 3, 4, 5, 6, 7, and 8 quarters of the inflation series. The contemporaneous correlation between the growth rate of monetary variables and inflation is presented when q=0.

Dynamic cross-correlations between the growth rate of selected monetary variables (*t*) and HICP inflation (*t*+*k*)
(first differences of the seasonally adjusted series)

	(q-8)	(q-7)	(q-6)	(q-5)	(q-4)	(q-3)	(q-2)	(q-1)	q	(q+1)	(q+2)	(q+3)	(q+4)	(q+5)	(q+6)	(q+7)	(q+8)
CC	-0.23	-0.13	-0.20	-0.02	-0.09	0.07	0.14	0.12	0.45**	0.15	0.18	0.26**	0.20	0.01	-0.07	0.14	0.08
M1	-0.22	-0.14	-0.25	-0.10	-0.05	-0.16	0.14	-0.04	0.22	0.23	0.22	0.37**	0.27**	0.28**	-0.03	0.04	0.09
M2	-0.22	-0.11	-0.21	-0.04	0.06	0.03	0.25**	0.18	0.37**	0.42**	0.25**	0.30**	0.15	0.18	-0.05	-0.08	-0.01

*Note: ** indicates statistical significance at 5% level.*

The table shows that the pattern observed in the data is the following: there is statistically significant contemporaneous relationship between money supply growth and inflation when CC and M2 are used. Moreover, inflation appears to act both as leading indicator to money supply (for M2) and as a variable that is lagging money supply dynamics for all money aggregates examined. When performing the same dynamic cross-correlations with the core consumer price index, the pattern is somewhat similar, however there seems to be much more evidence of a two-way relationship between developments in money and inflation. On the whole, when core inflation is examined, money-price relationships are somewhat less clearly defined than those for headline inflation.

Dynamic cross-correlations between the growth rate of selected monetary variables (*t*) and HICP core inflation (*t*+*k*)
(first differences of the seasonally adjusted series)

	(q-8)	(q-7)	(q-6)	(q-5)	(q-4)	(q-3)	(q-2)	(q-1)	q	(q+1)	(q+2)	(q+3)	(q+4)	(q+5)	(q+6)	(q+7)	(q+8)
CC	-0.38**	-0.26**	-0.33**	-0.20	-0.38**	-0.14	-0.03	0.09	0.26**	0.21	0.30	0.26**	0.24	0.25**	0.14	0.11	0.12
M1	-0.45**	-0.26**	-0.37**	-0.31**	-0.39**	-0.38**	-0.17	-0.16	0.07	0.13	0.28**	0.37**	0.37**	0.43**	0.31**	0.20	0.28**
M2	-0.30**	-0.19	-0.26**	-0.21	-0.19	-0.17	0.00	0.04	0.21**	0.31**	0.40**	0.43**	0.31**	0.35**	0.23	0.06	0.14

Note: ** indicates statistical significance at 5% level.

A serious drawback of dynamic cross-correlations that is generally not addressed in the empirical literature is the possibility for a spurious effect on the estimated correlation coefficients in the case of autocorrelation in the input series. As Haugh (1976) and Haugh and Box (1977) note in their research on checking the independence of two covariance stationary time series, the autocorrelation present in each of the series can inflate the variance of cross-correlation estimates above that expected when cross-correlating two white-noise series. For overcoming the problem of obtaining a likely distorted cross-correlation function with misleading inference on the pattern of cross-correlations, Haugh (1976) suggests first obtaining the appropriate univariate (e.g. ARMA) models for each of the series and afterwards estimating the cross-correlation function of the residual white noise series obtained by fitting each of the separate univariate models.

We apply Haugh's methodology by first pre-whitening money supply growth and inflation data, implementing the Box-Jenkins methodology[3] to the first differences of the seasonally adjusted series, and then proceeding with the automatic calculation of the sample cross-correlation function between the residual white-noise series. We identify Box-Jenkins models by following the standard guidelines for the behaviour of the sample

[3] We apply the Box-Jenkins methodology as detailed in 'Forecasting, Time Series and Regression: An Applied Approach' by Bowerman, B., R. O'Connel and A. Koehler, Fourth Edition, 2005.

autocorrelation function and the sample partial autocorrelation function. The appropriate univariate models are constructed and the necessary diagnostic checks to ensure that the residuals behave as white-noise processes are performed (see Appendix B and Appendix C).

We examine the Ljung-Box (1978) Q-statistic for high-order serial correlation of the residuals from the fitted models. Since the degrees of freedom are reduced by the number of estimated coefficients, we do an adjustment of the degrees of freedom on the modified Q-statistic for small samples.[4] The diagnostic tests on the residuals of the fitted models show that for all models the estimated statistic, estimated for 8 lags, is not statistically significant at the 5% significance level, *i.e.* the null hypothesis of zero autocorrelation in the residuals cannot be rejected.

We proceed with performing a cross-correlation of the residual series obtained from the estimated models. The residual cross-correlograms, given respectively for 8 lags and 8 leads, are presented below. The significance of each of the reported residual cross-correlation estimates is determined in the standard way by 2 standard deviations bounds of $\pm 2/(\sqrt{(T)})$, where T is the number of observations.[5]

Estimated univariate residual cross-correlation function for the growth rate of selected monetary variables (*t*) and HICP inflation (*t+k*)

	(q-8)	(q-7)	(q-6)	(q-5)	(q-4)	(q-3)	(q-2)	(q-1)	q	(q+1)	(q+2)	(q+3)	(q+4)	(q+5)	(q+6)	(q+7)	(q+8)
CC	-0.02	-0.22	-0.22	0.23	-0.16	-0.16	0.08	-0.09	0.37**	0.05	0.18	0.20	0.06	0.14	-0.22	-0.13	0.17
M1	-0.08	-0.11	-0.19	0.01	-0.14	-0.17	0.07	-0.18	0.22	0.01	0.12	0.28**	0.28**	0.18	-0.05	0.06	0.12
M2	-0.21	-0.05	-0.04	-0.10	0.10	-0.26	0.10	-0.11	0.22	0.14	-0.05	0.50**	0.01	0.18	0.08	0.11	0.03

Estimated univariate residual cross-correlation function for the growth rate of selected monetary variables (*t*) and HICP core inflation (*t+k*)

	(q-8)	(q-7)	(q-6)	(q-5)	(q-4)	(q-3)	(q-2)	(q-1)	(q)	(q+1)	(q+2)	(q+3)	(q+4)	(q+5)	(q+6)	(q+7)	(q+8)
CC	-0.04	-0.23	-0.15	-0.04	-0.05	0.06	0.02	0.18	0.00	0.32**	0.15	0.04	0.03	0.22	-0.16	-0.15	0.15
M1	-0.23	-0.14	0.00	-0.19	-0.21	0.11	-0.06	0.08	0.03	0.12	0.23	0.30**	0.22	0.31**	-0.11	-0.13	0.19
M2	-0.22	-0.04	0.03	-0.04	-0.05	0.04	-0.11	0.11	0.28**	0.21	0.09	0.14	0.20	0.25	0.03	-0.20	0.28**

*Note: ** indicates statistical significance at 5% level.*

[4] This modification in the statistic, which is used to test whether a chosen set of autocorrelations (e.g. for 8 quarters) in the residuals is significantly different from zero, is dependent on whether a constant has been included and as well as on the number of estimated autoregressive and moving average terms in the respective models. In each case, the value of the estimated modified Q-statistic is compared with the corresponding critical value from the chi-square distribution.

[5] As Haugh (1976) shows, the known asymptotic distribution of lagged cross-correlation estimates between two independent white noise series also applies to lagged cross-correlation estimates between the two residual series obtained by fitting univariate ARMA models to each of the two uncorrelated series.

We obtain the following results. For overall inflation and its relationship with the growth rate of currency in circulation, we observe a contemporaneous correlation. As regards the relationship between M1 and HICP inflation and M2 and HICP inflation, monetary aggregates lead inflation by about 3–4 quarters. Moving to the link between the selected monetary aggregates and core inflation, it can be seen that M2 has a significant contemporaneous relationship with core inflation and all monetary variables appear to have leading properties as regards inflation up to 8 quarters ahead. Furthermore, it does not appear that there is any feedback from inflation to monetary growth.

In addition to just comparing each cross-correlation coefficient with its respective standard error, we also apply Haugh's s statistic (1976) which presents a formal test of the null hypothesis that money supply dynamics and inflation series are unrelated, considering all examined cross-correlations. This statistic is computed as follows:

$$s = n \sum_{k=-M}^{M} r_{12}(k)^2 \text{ (for large samples);}$$

$$s^* = n^2 \sum_{k=-M}^{M} (n - |k|)^{-1} r_{12}(k)^2 \text{(for small samples)}$$

where: n is the number of time observations; $-M$ is the number of lags and M is the number of leads (M is chosen arbitrarily); $k = \pm\, 0,1,2,3...M$; r_{12} is the cross-correlation estimate at the specific lag/lead between the two residual series of length n, or when $k = 0$ (in case of the contemporaneous period).

In essence, the statistic is estimated for the first M residual cross-correlations and therefore M has to be determined by the researcher.[6] We estimate Haugh's s statistic for a different numbers of lags and leads of the cross-correlation function between the residuals from the fitted ARMA models to the series. The results, summarized in Appendix D, reveal a statistically significant feedback relationship between CC growth and HICP inflation, M1 growth rate and HICP core inflation, M2 growth and HICP inflation, and M2 growth and HICP core inflation. Furthermore, it can be seen that there is strong evidence for M1 growth leading core inflation and M2 growth leading overall inflation as well as core inflation.

[6] The critical value for this statistic is derived from the chi-square distribution (with $2M+1$ degrees of freedom for the case when k is from –M to +M and with M degrees of freedom when k is from 1 to M) under the null hypothesis that the series are uncorrelated at all lags considered.

The preliminary results presented in this section point to the conclusion that estimation methods based on the unidirectional assumption between money growth and inflation will probably be associated with inappropriate or even erroneous inferences. This is because when single equations are used for examining the association between the two variables, the OLS estimator will produce equations that are not identified correctly and precisely. We continue the empirical investigation by application of the standard Granger-causality tests to the original series of monetary aggregates growth and inflation.

3.2. Granger-causality Tests

This section is dedicated to the results obtained by an extensive Granger-causality test procedure using the OLS estimator in an unrestricted VAR framework. As indicated by the unit root tests, most of the series contain structural breaks, including in their first difference form. For this reason, we employ these tests with a rolling sample to tackle the problem of possibly instable coefficients, but also because we assume that structural changes that have likely taken place over the period from the late 1990s to 2012 can affect the relationship between the dynamics of a specific monetary aggregate and inflation throughout time. Therefore, that relationship and the possible existence of Granger causality can be different for different time periods. Two separate approaches are taken.[7] One is to use a sample with a rolling start, meaning that the end point of the sample is fixed for all tests (2012Q1), but the starting point is gradually moving forward in time. The other approach is to use a moving window sample with fixed length of eight years, meaning that both the start and endpoint of the sample are gradually moved forward.[8] Inflation and monetary variables are expressed in first differences of the levels (seasonally adjusted) in logs so as to achieve stationarity.

A critical step is determining the lag structure of each tested VAR system. On the one hand, choosing too few lags can result in omitted variable bias, which will produce a specification error and the estimation of incorrectly specified coefficients. On the other hand, using too many lags is

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[7] For a description of possible rolling methods applied in the empirical literature, see for example Levasseur, S. (2008). The author provides definitions of the following techniques: 1) 'pure rolling method' – an arbitrarily chosen length of window (a fixed number of quarters) is rolled over the sample under investigation; 2) 'recursive method' – the starting date is kept fixed and the window size grows as the end date is approached; 3) 'reverse recursive method' – the end date is fixed and the window size shrinks as the distance from the starting date increases. Here we apply the 'reverse recursive method' and the 'pure rolling method'.

[8] Eight years is picked with the thought of having at least around 30 observations for each test.

wasting valuable observations and such specification error can result in the estimation of unreliable coefficients with high standard errors. Since the lag structure is crucial for the correct estimation of the tested system, the Akaike Information Criterion (AIC) and the Schwartz Criterion (SC) were consulted. However, the number of lags in some cases was determined subjectively, irrespective of the results from the two information criteria, in order to avoid autocorrelation in the residuals. Dummy variables were also used subjectively for each set of variables in order to fulfill the criterion of normality in the distribution of residuals.

The testing procedure involved estimating over 200 different unrestricted VAR models and a summary of the results is presented in Appendix E. For the first sampling method of a moving start of the sample period several tendencies are observed: for different sub-samples starting between 2000Q1 and 2002Q1 and ending in 2012Q1, there are statistically significant results pointing out to causality running from CC to headline inflation; however, the opposite causality is observed in some of the sub-samples as well; for sub-samples starting from 2000Q4 to 2004Q4, with a few exceptions, the results show causality running from M1 to headline inflation; testing for causality between M2 and headline inflation yields very mixed results; tests for CC and core inflation are mostly resulting in causality running from the former to the latter; unambiguous results also point out to causality running from M1 to core inflation almost throughout the whole period; testing M2 and core inflation, again, produces mixed results, but they mostly point out to either causality running from M2 to core inflation or to two-way causality.[9]

For the second method with a moving sample window the findings can be summarized as follows: testing headline inflation and CC yields mostly inconclusive results; significant results point out to causality running from headline inflation to M1 for the sub-samples 1998Q1–2006Q1, 1999Q1–2007Q1 and 2000Q1–2008Q1; for three other later sub-samples causality goes from M1 to overall inflation; for the sub-sample 1998Q1–2006Q1 results point out to causality running from headline inflation to M2, however for two other later sub-samples we find causality from M2 to headline inflation; testing core inflation and CC shows that for three sub-samples currency in circulation granger causes core inflation; significant results point out that causality runs from M1 to core inflation for

[9] When we use the word 'causality' for short, we actually mean Granger causality and we do not refer to the existence of an actual cause and effect relationship, that is we refer to the traditional Granger causality understanding in the sense of one variable preceding another variable in time.

four of the seven sub-samples; results from tests of M2 and core inflation yield inconclusive results and no causality in most of the sub-samples.

It is evident that the two methods have some differences in their findings. However, a few common results can be outlined: both sampling methods yield test results on currency in circulation and core inflation pointing out to causality running from the former to the latter; causality from the monetary aggregate to core inflation as well as to headline inflation is found for tests on M1, with this causality being stronger for the case of core inflation; tests involving M2 yield very mixed results whether headline or core inflation is used and often the tests show two-way causality. It should be also noted that sub-samples starting more recently generally show causality running from the monetary aggregates to inflation, while the opposite is true for sub-samples starting earlier. Therefore, our testing procedure proves that granger causality exists and in the predominant number of cases causality goes from monetary growth to inflation. However, the nature of that causality changes throughout time and this result may be probably attributed to the structural breaks that were revealed for most of the series in the unit root tests. Furthermore, significant evidence of two-way causality between M2 and inflation is also found.

3.3. Issues for Robustness of Granger-causality Tests

Formulating conclusions based on the Granger-causality test results described above needs to be necessarily accompanied with an appropriate assessment of possible cointegration. In the presence of the latter, the inclusion of 'error-correction terms' in Granger-causality tests is required so as to avoid misspecification due to an omitted variable bias problem related to cointegration. There are two possible econometric options for dealing with potential cointegration between monetary aggregates and prices/inflation. The first one is to re-estimate Granger-causality tests, including where necessary co-integration links. The second option is to apply the system approach framework of Johansen (1990, 1991) which enables testing for more than one cointegrating relationship in a vector autoregressive model, possibly including not only money and prices, but also all standard variables related for example to money demand investigation in a multivariate analysis. One of the main weaknesses of the first option is that even if one omitted variable bias problem can be addressed by adding cointegrating term for the two variables, there may be other potential serious problems causing misspecification such as the issue related to omitting relevant explanatory variables in the cointegration relation or omitting whole cointegrating vectors.

Our preference for applying the Johansen procedure is driven by the following main reasons. First, this procedure is the natural way of incorporating long-run relations, one of them in the case of money and prices/inflation is from a theoretical point of view expected to be naturally related to the estimation of a full money demand function. The adjustment of monetary aggregates to the long-run money demand function can be interpreted as money supply reaction to money demand, provided the estimated model is correctly specified with all relevant variables for a money demand function. Furthermore, possible adjustment of price variables can be thought of as the case when any short-run disequilibrium between money supply and demand results in excess money supply potentially acting as a pro-inflationary factor which represents the second hypothesis postulated to be tested in the paper. In addition, the short-run dynamics of the VECM allows for testing the impact of past values of monetary variables growth on inflation (as well as the impact of past values of inflation on monetary variables growth) once any feedback from disequilibrium in the cointegrating relationship has been taken into account. This can also help investigate further the first hypothesis of the paper.

4. Estimates of Money Demand Functions with VECM

4.1. Brief Overview of the Theoretical Framework

The general theoretical framework for modeling money demand in a specification relevant for the long-run horizon can be summarized as given below:

$$\frac{M}{P} = f\,(Y, R) \qquad (1)$$

where M represents the monetary aggregate in nominal terms; P is the price level measured usually either by the consumer price index (HICP/CPI) or the GDP deflator; Y is the scale variable characterizing the transaction/precautionary motive of holding money and could be represented by variables such as real consumption expenditures of households, real wages or real retail sales (when demand for currency in circulation is modeled) and real GDP or industrial production (when broader monetary aggregates are modeled); R represents a vector of returns, including *own rate* of money which is to be chosen depending on the composition of the specific monetary aggregate as well as different indicators that proxy the opportunity cost of holding money for the respective monetary aggregate such as the return on assets *outside of money* and the (expected) rate of inflation, with the latter representing the opportunity cost of holding money rather than goods or real assets in general. According to the standard theories of money demand, the function f is increasing in Y, increasing in the components of R that relate to assets included in M and respectively decreasing in the components of R that relate to assets outside of M.

In the cointegrated VAR modeling that we are going to employ we treat prices as an endogenous element to the system of analyzed variables and use (1) in the following form:

$$M = f\,(P, Y, R) \tag{2}$$

The endogenous treatment of the price level allows us to test whether the theoretical hypothesis for long-run unit price elasticity is supported by the data without imposing this restriction beforehand as in (1). We thus depart from the general specification structure of most empirical studies which assume long-run price homogeneity and additionally assume the validity of the hypothesis that money supply is not an important factor affecting price dynamics, thus justifying the omission of prices from the vector of endogenous variables in (1). By including the price level among the set of endogenous variables and attempting to model nominal rather than real money demand, our approach broadly resembles that of Komárek and Melecký (2001) who construct a money demand model for the Czech Republic. The characteristic feature of their paper, presented in the literature review section, is the importance that the authors attach to transition-specific factors and one example of such factors is the preference for modeling money demand in nominal terms.[10]

In a log-linear form the model we estimate for CC, M1 and M2 is as follows:

$$m^d = \gamma_0 + \gamma_1 y + \gamma_2 R^{\text{out}} + \gamma_3 R^{\text{own}} + (\gamma_4 \Delta p) + \gamma_5 p$$

Expected signs: $\gamma_1 = 1$, $\gamma_2 \leq 0$, $\gamma_3 \geq 0$, $\gamma_4 \leq 0$, $\gamma_5 = 1$

4.2. Description of the Data and Methodology

This part of the paper reports the results of money demand models for CC, M1 and M2 obtained by applying the Johansen cointegrated VAR approach. We investigate a large dataset for the analysis of the three

[10] The authors acknowledge that for transition economies there is uncertainty with respect to the specific price variable to be used as a deflator for estimating real money holdings due to factors such as the severe shocks related to deregulation of prices, transformation of tax policies, liberalization of foreign trade, *etc.* that have occurred, affecting the specific evolution of different measures of price developments over time.

measures of money, being aware of the fact that the specific choices of explanatory variables representing Y, P and R in (2) may significantly affect the outcome of the cointegration test for money demand relations.[11]

Based on the detailed literature review, we select the following possible variables as the dataset to be tested for modeling CC money demand: HICP (as the price variable); real expenditures of households, real wages, real retail sales (as proxies for real sector developments); and the BGN/USD exchange rate and the interest rate on new term deposits (as opportunity costs of holding currency in circulation). The dataset for modeling M1 covers a broader range of variables: HICP, GDP deflator and producer price index (PPI) (as possible price variables); real GDP, industrial production, real retail sales (for approximation of real income) and FDI as a variable reflecting the substantial amount of foreign direct investment in the Bulgarian economy with a potential influence on demand for M1; interest rates on overnight deposits (as a proxy for the own rate of money due to the relatively large share of overnight deposits in the composition of M1); the interest rate on new term deposits, the BGN/USD exchange rate and the 3-month money market rate SOFIBOR (as proxies for the opportunity cost of holding M1). The dataset for modeling the broad monetary aggregate includes: HICP, GDP deflator and PPI; real GDP, industrial production; the interest rate on new term deposits (as a proxy for the own rate of money); the interest rate on loans, the 3-month money market rate SOFIBOR, the 3-month EURIBOR, the BGN/USD exchange rate, the 10-year Bulgarian government bond yield and the 10-year euro area government bond yield (as possible proxies for the opportunity cost of demand for M2). Furthermore, for the specifications of all money aggregates, it is additionally investigated whether the inflation rate could enter as a long-run determinant of money demand in its role of an opportunity cost of holding money rather than goods or real assets.

A few notes need to be made regarding the selection of variables. The initial period of low monetization immediately following the introduction of the currency board arrangement, the gradual development of the banking sector, the stock market and the government bonds market, implying an initial period with practical absence of alternatives for storing wealth, the changes in the composition of monetary aggregates, the process of

[11] See Ericsson (1998) for a thorough discussion of the key issues in the empirical modeling of money demand, including data measurement and data choice problems, parameter constancy, the adequate construction of the opportunity cost of holding money as well as various specification issues that have to be considered. Ericsson argues that in testing for cointegration across different measures of money, scale variables, and rates of return, only a few combinations might be cointegrated, even if the underlying cointegration relationship exists.

financial innovation, the presence of grey economy are all reasons why it may be difficult to find appropriate indicators for opportunity costs of holding money. For this reason, we start off with an investigation of a relatively large dataset of possible explanatory variables in the money demand models. It is worth noting that among the potential measures for the opportunity cost of holding M2 along with traditional indicators in the economic literature we add the interest rate on loans as suggested by Komárek and Melecký (2001) who analyzed broad money demand in the Czech Republic and Dabušinskas (2005) who estimated a broad money demand model for Estonia.

The period of empirical analysis is from 1998Q1 to 2012Q1. However, for most of the models investigated the period starts from 2000Q1 and for some from 2003Q1 because of data availability. We work with seasonally adjusted quarterly data, with the adjustment performed in Eviews using the Tramo/Seats procedure with the automated setup for the diagnostics and optimization. The procedure does not identify seasonality for interest rates and the producer price index. The aggregation for the data which is published on a monthly frequency is obtained by averaging. The only exception are the series of the monetary aggregates, for which as explained in Section 3, we take an end-of-period transformation.

Estimations of money demand models within the Johansen framework are conducted without testing the order of integration of the variables as this is found not to be a necessary prerequisite for the analysis.[12] We have identified in Section 3 that monetary aggregates as well as the HICP are $I(1)$ processes in levels and in addition, unit root tests applied to the other two price indicators employed in the analysis (the GDP deflator and the PPI) show that these series are also integrated of order one (see Appendix A1 and A2). Consequently, no exhaustive examination of the stochastic time-series properties of the rest of the variables is performed.

The initial VAR models that are tested include levels of the variables expressed in logs with only interest rates expressed in percentages. A few variables such as the BGN/USD exchange rate, the 3-month EURIBOR, the 3-month SOFIBOR and the FDI stock are treated as exogenous both because such an assumption appears plausible for most of them, but also because treating them as endogenous usually produced very large systems to be estimated and made the application of the Johansen method unreliable due to the relatively small dataset. The inclusion of the inflation rate

[12] As Johansen (1995), cited in Ahking (2002), argues in case where there are at least two non-stationary variables in the cointegration model that are integrated of the same order, the other variables could be both stationary and/or trend-stationary.

in the VAR models is achieved either by considering the annual rate of inflation or by taking lags of the first difference of the log-level of the harmonized consumer price index.

The selection of the lag length is done in accordance with the information criteria (maximum lag is set to 4), with strongest reliance given to the Schwarz criterion due to the small sample. The final decision on the lag length of the respective model is made once a well-specified model with serially-uncorrelated and normally-distributed residuals is obtained and the VAR displays dynamic stability. When normality of residuals is difficult to achieve, the minimum requirement is that the distribution of residuals satisfies the assumption of non-skewness.[13] We also implement the adjustment of the critical values of the trace test as a correction for finite sample sizes, as proposed by Reinsel and Ahn (1988) and cited in Cheung and Lai (1993).

As it is known, the power of the Johansen cointegration test may weaken when exogenous series are added. We therefore choose the specific number of cointegrating vectors for each model investigated based on the trace test in combination with the economic interpretability of results. Once the cointegrating rank is determined, the specific number of cointegrating vectors is imposed on the VAR system based on the general assumption that there are linear trends in the data, while the cointegration relation(s) contain a non-zero intercept but no deterministic trends. Estimation procedure then continues with imposing restrictions on the cointegrating vector(s) so as to ensure their unique determination. Tests for significance and stationarity of the variables in the cointegrating relation(s) are performed. Then restrictions based on the postulations of money demand theory for one of the cointegrating vectors are imposed and in case a second vector is determined, other restrictions such as those related to the term structure between interest rates or to aggregate demand relations have been tested. Weak exogeneity of some of the variables is also investigated. As a final step, specification tests are performed.

4.3. Results for Currency in Circulation

The empirical analysis of the currency in circulation established a VECM for the following variables: currency in circulation (cc), HICP (hicp), real retail sales[14] (tradec) and the interest rate on new term deposits (m_tdir_n). The

[13] We follow the recommendations by Hendry and Juselius (2000) who show that statistical inferences are valid even if there is empirical evidence for excess kurtosis and residual heteroscedasticity. The authors also suggest using the trace test for determining the number of cointegrating vectors as a more robust indicator than the maximum eigenvalue test.

[14] Total turnover index of retail sales, except of motor vehicles and motorcycles at constant prices

variables are seasonally adjusted and expressed in logarithm, apart from the interest rate variable. The inclusion of the inflation rate as an additional opportunity cost produced a model with unstable coefficients.

The unrestricted VAR passed successfully specification tests for residual autocorrelation, normality of residuals and heteroscedasticity when the model included two lags as well as two dummy variables.[15] The test for the VAR stability condition check pointed out to dynamic stability of the investigated model. The Johansen cointegration trace test results, given in Appendix F1 (Section 1), reveal an empirical evidence for 2 cointegrating relationships. We thus proceed with imposing rank = 2 on the unrestricted model and make the assumption that there are linear trends in the data series with no deterministic trends in any cointegrating relations.[16] We succeed in obtaining unique specifications for these cointegrating relationships with the following over-identifying restrictions which are jointly accepted (LR test of restrictions: Chi^2(2) = 0.484 [0.7850])[17]:

cointegrating vectors	β_1	β_2
m_tdir_n	0.073 (0.009)	0.112 (0.018)
log(cc_sa)	1.0	0
log(hicp_sa)	-1.0	1.0
log(tradec_sa)	-0.885 (0.019)	-1.0

adjustment coefficients	α_1	α_2
m_tdir_n	-1.179 (0.883)	-1.127 (0.305)
log(cc_sa)	-0.507 (0.086)	0
log(hicp_sa)	-0.183 (0.045)	-0.089 (0.017)
log(tradec_sa)	-0.119 (0.062)	-0.065 (0.024)

Note: standard errors in parentheses.

The first cointegration relationship, which can be seen in Column 2 of the table with the beta coefficients, can be interpreted as a money demand function. In the long run currency in circulation is driven by the increase in the price level (with accepted long-run price homogeneity) and in the retail sales volumes (coefficient of 0.885, slightly less than the unit elasticity as postulated by the quantity theory of money) and is influenced negatively by the interest rate on new term deposits (the semi-elasticity stands at -0.073). These results are in line with theoretical considerations. The generally low interest rate elasticity also corresponds to empirical results for EU New Member States. The loading coefficient for the *cc equation,* which is presented in the second column of the table with the alpha coefficients,

[15] For the fourth quarter of 2000 and the fourth quarter of 2008 respectively.

[16] We tested the hypothesis whether a deterministic trend can be included in the cointegration relations, but that hypothesis was rejected.

[17] The asymptotic null distribution of the test is denoted Chi^2(·) with degrees of freedom in parentheses and the asymptotic p-value given in the brackets.

is negative at -0.51 which points to relatively fast adjustment of *cc* to short-run deviations from long-run equilibrium. This corresponds to expectations as under the currency board demand and supply of currency in circulation should be balanced in a fast and automatic way. The loading coefficient for the *hicp* is negative at -0.18 which does not support the hypothesis that in cases of an excessive supply of currency in circulation inflation increases.

The second cointegration relationship is set for the *hicp* (the normalization is done on this variable). In the long run *hicp* is related positively to the real retail sales (a permanent increase in real retail sales by 1% increases the consumer price index by 1%) and negatively to the interest rate on new term deposits (an increase of interest rates by 1 percentage point decreases the hicp index by 0.11%). The loading coefficient for the *hicp* equation in this cointegration relationship is as expected negative and shows relatively slow adjustment of prices to disequilibrium.

The results from the recursively estimated LR test of restrictions, presented in Appendix F1 (Section 2), show that the imposed restrictions are valid for the whole period of estimation which attests to the stability of the results. In the addition, the recursively estimated unrestricted coefficients (on the interest rate and on the real retail sales in the first cointegration equation and on the interest rate in the second cointegration equation) are also stable (Section 3).

We then transform the model in a stationary form (that is into a model with the respective error-correction terms combined with short-run dynamics of the variables), being particularly interested whether the short-run dynamics of the *hicp* variable is influenced by the dynamics of the *cc* variable. We obtain a statistically significant positive coefficient of 0.21, showing that while the dynamics of *cc* is affected by the development of prices in the long run, there is also a feedback influence from *cc* to the dynamics of prices (see Appendix F1, Section 4). Furthermore, the equations from the short-run error correction model successfully pass all residual diagnostic tests both when run for the case of single equations and for the vector version of the tests (Section 5). The recursively estimated 1-step residuals of the model for each of the four variables display no problems with instability (Section 6).

4.4. Results for M1

For the monetary aggregate M1 we obtain an empirical evidence for a cointegrating relationship with the following variables: producer price index (ppi), industrial production index (indp) and the interest rate on new term deposits (m_tdir_n). M1 and industrial production are seasonally adjusted and expressed in logarithm, no seasonal adjustment was required for the producer price index and this index is only expressed in logarithm, whereas for the interest rate variable no transformation was applied.

The cointegration of *m1* with *ppi* as the price variable can be accounted for by the fact that a predominant part of the composition of M1 consists of overnight deposits that are hold mostly by non-financial corporations rather than households over the period under investigation. An additional money demand determinant that was also explored was the exchange rate BGN/USD (*er)* which could potentially influence M1 holdings for example in the case when a depreciation of the national currency against the US dollar could theoretically be associated with increasing external demand for Bulgarian goods, respectively higher domestic industrial production, higher domestic inflation rate and a need for more money in the economy because of the larger amount of transactions.

We obtain a well-specified model with a VAR (1) when including as exogenous variables the second lag of the log-level of *er (log(er(-2))*, the first difference of the log-level of the index of Brent oil prices measured in euro (*oil_p*): *dlog(oil_p)* (to account for residual autocorrelation in the *ppi* equation) and two dummy variables[18]. This model produced the required assumptions for normality of residuals, lack of serial autocorrelation as well as for dynamic stability. In addition, the adequacy of the chosen model was supported by the lack of residual heteroscedasticity both for the single equations and for the overall model (based on the vector heteroscedasticity test). The application of the Johansen cointegration trace test and the small-sample correction of the trace statistic, as discussed in Section 4.2, provided evidence for one cointegrating vector (see Appendix F2, Section 1).

Imposing one cointegrating vector in the unrestricted VAR model and working with the assumption about linear trends in the data, without any linear trends in the long-run relation, we proceed by testing hypotheses about the statistical significance of the variables in the long-run relation.[19] The results from the tests (not reported here for reasons of brevity) showed that the coefficient on the interest rate is not significantly different from zero. This fact is in contrast with theoretical postulates about a negative coefficient as the interest rate on new term deposits represents an opportunity cost of holding money in the form of the two subcomponents of M1. For the purpose of ensuring an identified model which can be estimated, we start

[18] The dummy variables are for the third quarter of 2008 and the first quarter of 2009 respectively.

[19] We also estimated the same model with one cointegrating vector based on the assumption that a restricted trend lies in the cointegration space, but in this model the trend was not statistically significant.

with the hypothesis of an existing money demand function and impose the standard theoretical restriction of unit price elasticity in addition to the restriction of zero interest rate elasticity. While these two restrictions are not rejected, we explore the statistical significance of the loading coefficients in the *ppi* and *indp* equations and subsequently apply two other restrictions – for weak exogeneity of *ppi* and *indp*. The joint LR test for all these over-identifying restrictions shows that the restrictions cannot be rejected (LR test of restrictions: Chi^2(4) = 6.119 [0.1904]). The obtained cointegrated VAR model with these restrictions is presented below:

cointegrating vector	β_1
m_tdir_n	0
log(m1_sa)	1.0
log(ppi)	-1.0
log(indp_sa)	-2.385 (0.189)

adjustment coefficients	α_1
m_tdir_n	-1.206 (0.261)
log(m1_sa)	-0.159 (0.040)
log(ppi)	0.0
log(indp_sa)	0.0

Note: standard errors in parentheses.

As can be seen, in the long-run income elasticity exceeds one (2.385), which is similar to the results generally obtained for EU New Member States. The established cointegrating vector can be interpreted as a money demand function, which is also supported by the negative adjustment coefficient in the *m1* equation (-0.159). The interest rate on new term deposits adjusts to deviations from the long-run relation with a negative coefficient, implying that any excess money supply brings about a decrease in the interest rate. Based on the weak exogeneity of *ppi* in this model and the lack of effects from lagged differences of *m1* on *ppi* dynamics (due to the lack of first-differenced terms in the vector-error correction model), we can conclude that there is no feedback effect from *m1* to *ppi*.

The results from the recursively estimated LR test for the four restrictions discussed above, given in Appendix F2, Section 2, contest to the overall stability of restrictions with the exception of the period around the end of 2009 and the beginning of 2010. For that period we can tentatively argue that the downturn of the Bulgarian economy might have resulted in changes in the behavior of economic agents and this could potentially destroy the validity of the restrictions on the model. At the same time, however, the recursively estimated unrestricted coefficient for the income elasticity displays overall stability as shown in Appendix F2, Section 3.

Transforming the VECM model into a stationary short-run form, we note additional results. The rate of increase in the price of oil has a statistically significant positive influence on the *ppi* index, which can be accounted for by the relatively large weight of energy in the producer price index. The loading coefficient for *ppi* is negative (at -0.024) which does not lend support to the hypothesis of a pro-inflationary impact of excess supply of M1 (Appendix F2, Section 4). In Section 5 we report all required residual diagnostic tests and it can be seen that all test results are in general satisfactory.

4.5. Results for M2

Econometric modeling for a cointegrating relationship including M2 led to empirical evidence for a model including M2, the producer price index (*ppi*), the industrial production index (*indp*), the interest rate on new term deposits (*m_tdir_n*) and the interest rate on loans for new business (*ir_loans*). The endogenous variables were taken in logs and seasonally adjusted, apart from the two interest rates variables. No seasonal adjustment was required for the *ppi*.

To obtain a well-specified model we considered a number of alternatives for the lag structure as well as a number of dummy variables and additional exogenous variables to correct mostly for normality of residuals (targeting at the least at their non-skewness), as lack of residual serial correlation was hard to attain. The dynamic stability properties of the initial unrestricted VAR were also a requirement that had to be satisfied. The final specification included 2 lags of the endogenous variables and as exogenous to the system the following variables: the first lag of the exchange rate BGN/USD *(log(er(-1))*, the first difference of the logarithm of Brent crude oil price (in euro): *dlog(oil_p)*, the first difference of the logarithm of real GDP in the euro area (seasonally adjusted): *dlog(rgdp_ea_sa)* as well as dummy variables.[20] This specification produced a statistically adequate unrestricted VAR model.

For this model the Johansen cointegration trace test revealed the presence of 1 cointegrating vector (see Appendix F3, Section 1). Obtaining unique specification for the cointegrating relationship is achieved with the following over-identifying restrictions which are jointly accepted (LR test of restrictions: Chi^2(3) = 0.796 [0.8504]):

[20] The following dummy variables were included: for Q1 2005, Q2 2005, Q1 2007, Q4 2008 and Q1 2009.

cointegrating vector	β_1
m_tdir_n	-0.072 (0.011)
ir_loans	0.068 (0.012)
log(indp_sa)	-1.0
log(m2_sa)	1.0
log(ppi)	-1.454 (0.093)

adjustment coefficients	α_1
m_tdir_n	-1.026 (0.711)
ir_loans	0
log(indp_sa)	0.177 (0.075)
log(m2_sa)	-0.165 (0.038)
log(ppi)	0

Note: standard errors in parentheses.

The imposed restriction in the cointegrating relationship follows the conventional money demand models and is related to unit income elasticity. Additionally, the adjustment coefficients in the equations for *ir_loans* and for *ppi* are set to zero. For the remaining unrestricted coefficients, we note that *m_tdir_n* enters the long-run relationship as expected with a positive sign and *ir_loans* has the expected negative coefficient. Long-run price elasticity comes to 1.454. As regards the adjustment coefficients, we can observe the expected negative coefficient for the M2 equation. The industrial production index responds positively to deviation of M2 supply from the relationship with the long-run components of money demand, meaning that 'excess money supply' affects economic performance. The results from the recursively estimated LR test for the three restrictions show stability properties (Appendix F3, Section 2). The recursively estimated unrestricted coefficients in the cointegrating relationship also display overall stability properties with a slight instability observed at the end of 2009 (see Section 3).

Transforming the cointegrated VAR model to a system with $I(0)$ variables leads us to the model equations presented in Section 4. In the short-run equation for the *ppi* the previous period growth of M2 has a statistically significant positive effect. The equations from the short-run error correction models pass successfully almost all residual diagnostic tests: both when run for each of the single equations and as in the vector version form (see Appendix F3, Section 5). The only tests that are not satisfied are the test for residual autocorrelation in the industrial production equation, the test for heteroscedasticity in the equation for the interest rate on new term deposits as well as the vector tests for normality and heteroscedasticity. The recursively estimated 1-step residuals from the five short-run equations are constant as shown in the graphs in Section 6.

5. Conclusions

This paper has undertaken an extensive empirical analysis of the relationship between money supply dynamics and inflation, continuing and further deepening the research conducted in the first years of the introduction of the currency board arrangement in Bulgaria.

The first hypothesis tested in the paper was whether a feedback relationship exists from money aggregates dynamics on inflation. The results based on univariate residual cross-correlation approach point out to the fact that any estimation methods based only on the unidirectional assumption (from prices to money supply) are most probably associated with inappropriate or even erroneous inferences. The application of Granger-causality tests shows that granger causality exists in both directions and in the predominant number of cases causality goes from monetary growth to inflation. There is evidence that a one-way direction of relationship exists from M1 growth to HICP inflation and this relationship is particularly strong from M1 growth to HICP core inflation. We also obtain results on currency in circulation and HICP core inflation that generally point out to causality running from the former to the latter, with this evidence being somewhat weaker when currency in circulation and headline inflation are analyzed.

Apart from the new conclusions from empirical econometric analysis on the relationship between money supply dynamics and inflation, an additional important contribution of the paper includes the results from cointegrated VARs related to the estimation of money demand models for currency in circulation, M1 and M2. Our main conclusions about money demand models can be summarized as follows.

We obtain a stable long-run money demand relation for currency in circulation, HICP, retail trade turnover at constant prices and the interest rate on new term deposits, with all coefficients having the theoretically expected signs and unit price elasticity being confirmed. The adjustment of currency in circulation to deviations from long-run equilibrium is relatively fast (adjustment coefficient at around -0.5) which is in line with expectations as under the currency board arrangement demand and supply of currency in circulation should be balanced in a fast and automatic way. The short-run dynamics reveals feedback influence from money to price dynamics. The overall parameter constancy tests of the model do not show any problems regarding changes in the sample period.

Concerning M1 and M2, cointegrating money demand relationships were more difficult to obtain and the models explored remain with some specification problems especially with respect to stability of coefficients and the choice of variables to be included. We obtain evidence for a long-run link among M1, the producer price index, the industrial production index and the interest rate on new term deposits. In this relationship unit price elasticity is confirmed, whereas the income elasticity of money demand, in line with findings for other new EU Member States, is higher than 1 (2.4). The main issues with the model relate to the insignificance of the interest rate variable in the long-run link and the instability of the cointegrating relationship at the end of 2009 which needs to be further investigated. The model does not lend support for a possible feedback effect from the short-run dynamics of M1 back to that of producer prices.

For M2 we reach the conclusion of a long-run money demand relationship with the producer price index, the industrial production index, the interest rate on new term deposits and the interest rate on loans for new business. All of the coefficients have the theoretically expected signs, unit income elasticity of money demand is obtained, while the long-run price elasticity is slightly higher than 1 (1.4). In the short-run equation for the producer price index, the previous period growth of M2 has a positive effect on the dynamics of producer price inflation. Similar to the model for M1, this model displays certain instability at the end of 2009.

Finally, regarding the second hypothesis investigated in the paper whether excess money supply in the short-run might act as a pro-inflationary factor, the empirical analysis based on the cointegrated VAR models does not lend support for such an effect for any of the monetary variables analyzed.

To summarize, the results of the paper point to the general conclusion of a two-way relationship between money supply and price dynamics in Bulgaria. The link from prices through money demand to money supply may be theoretically justified because of the endogeneity of the money supply mechanism under the currency board arrangement. At the same time, the paper also finds evidence for feedback effects from monetary growth to inflation.

Future areas of research on the relationship between money supply dynamics and inflation include: 1) exploring in more detail the properties of M1 and currency in circulation for improving forecasts of core inflation; 2) the application of Johansen cointegration analysis with structural instability to deal with the identified structural breaks in most monetary aggregates and price variables series as well as with the evidence for stability problems of money demand functions for the cases of M1 and M2; 3) the extension of estimation techniques for money demand models with methods such as the single-equation techniques of DOLS and DGLS with the aim of providing evidence for the robustness of results.

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