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THE IMPACT OF MONEY ON OUTPUT IN CZECH REPUBLIC AND ROMANIA

Mihaela SIMIONESCU[1], Adam P. BALCERZAK[2], Yuriy BILAN[3*], Anna KOTÁSKOVÁ[4]

[1]Institute for Economic Forecasting, Bucharest, Romania
[2]Department of Economics, Nicolaus Copernicus University, Toruń, Poland
[3]Faculty of Management and Economics, Centre of Applied Economic Research, Tomas Bata University in Zlin, Zlín, Czech Republic
[4]Faculty of Economics and Business, Paneuropean University in Bratislava, Bratislava, Slovakia

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Abstract. The problem of relationship between output and money has become again a subject of special interests of economists after the most recent global financial crisis and monetary stabilization policies applied by central banks of almost all developed economies. In this context, the main aim of this paper is to assess the relation between GDP and the most important monetary variables in two countries: Romania and Czech Republic over the period of 1995:Q1 – 2015:Q4. The choice of these economies was deliberate. The selected countries are different from the viewpoint of rate and results of transformation from the centrally planned to market economy, which have influenced their current economic environment stability. Czech Republic is currently classified as middle or even developed country, whereas Romania is still considered as a developing economy. Thus, differences between these two countries make them interesting in the case of comparative studies. In the empirical part of our research the vector error correction models (VECM) were applied. The main findings of the article are the following: in Romania, there is a short-run causality from money supply (M3) to GDP and a long-run relationship between GDP, internal credit and M3. According to Granger causality test, the rate of M3 in Romania was a cause for economic. In Czech Republic, there is a short-run causality from M3 to GDP and a long-run causality between GDP, internal credit and M3. Thus, the results contradict the money neutrality hypothesis in post-transformation Central European economies.

Keywords: GDP, VECM, internal credit, money supply, money demand, neutrality of money, Granger causality.

JEL Classification: C11, C13, C51.

*Corresponding author. E-mail: *yuriy_bilan@yahoo.co.uk*

Introduction

The debate concerning the effectiveness of monetary policy and its influence on short- and long-term growth has been present in the core of modern macroeconomics since its beginning starting from the publication of The General Theory of Employment, Interest and Money. The importance and actuality of this topic is related to the construction of sustainable monetary policy, which is the condition for ensuring healthy economic growth. Thus, both academic economists and policymakers have been interested in measuring the potential influence of monetary variables on output. The issue has been the subject of intensive research in literature (starting with Sims 1972; Stock, Watson 1989; King, Watson 1997). Since the end of 1980 till the beginning of 2000s, the role of money was overlooked in a large number of studies (Taylor 1999; Clarida *et al.* 2000).

At the end of the previous century it seemed there was a commonly accepted consensus concerning the concept of middle- and long-term neutrality of money. There could be also seen a strong skepticism as to application of monetary tools for the objectives of short-term stabilization. It was commonly believed that central banks should only concentrate on the inflation stabilization purposes. From the policy guidelines perspective, it could be seen, for example, in the Maastricht treaty criteria or the suggested good governance practices related to first Washington consensus. However, monetary stabilization actions conducted in the United States after 2000 have contributed to the renewal of research interests in the relation between monetary policy tools and output (see Belongia, Ireland 2016). As a result, currently many economists suggest that money can be still a strong policy instrument during an economic crisis (Nelson 2003; Duca, Van Hoose 2004; Sims, Zha 2006; Hill 2007; Yıldırım 2015). The research topic has also gained special actuality after the latest global financial crisis and the related large-scale monetary stabilization policies applied by almost all developed and many developing economies as well (Janus 2016; Jędrzejowska-Schiffauer, Schiffauer 2016; Palankai 2015; Svitálková 2014).

In this context, the main aim of this paper is to analyze the relationship between output and monetary variables using some econometric approaches (Granger causality on stationary data, vector error correction models, dynamic panel data models). The research was conducted for Czech Republic and Romania for the years 1995:Q1 – 2015:Q4. We selected these countries deliberately based on their comparative value from the perspective of comparative studies in Central European post-transformation economies concerning the effectiveness of monetary policy tools in different institutional and macroeconomic environment (Lakic *et al.* 2015). The countries have had different paths in their economic development after 1990. Czech Republic succeeded in making quick transition to the market economy, while Romania still faces problems in getting a sustainable economic growth. Czech Republic is currently the most developed Central European economy; it is also a member of OECD, whereas Romania is still considered as a developing economy. Thus, this makes these two countries an interesting case study from the comparative perspective. It is important to see in these two types of economies if there are still differences regarding the money-output relationship.

This contribution can be placed into the current empirical research context concerning the concept of money neutrality in case of the economies being at different development levels. The key novelty of our research is the approach based on time series to study long-run and short-run relationship completed by a panel data approach based on dynamic panel models for the two post-transformation countries. Moreover, this study is not reduced to causality analysis, estimations being provided for each country.

In the next section, review of the previous empirical literature concerning highly developed and developing countries is provided. Then the methodology used in the empirical research is briefly described. The results of econometric estimations are presented in the next part, while the last section concludes.

1. Review of previous empirical research for highly developed and developing economies

New approaches from literature have currently focused on the money-output causality in the light of complex methods. Starting with the developed economies the money-output causality for the United States in the years 1959–2006 was verified by Gefang, who was using the logistic smooth transition vector error correction model (Gefang 2012). In the postwar period, a nonlinear relationship between money and output was detected, the changes being due to price levels and increases in output. In the traditional approach, money was not a Granger cause for output. Only a nonlinear causality between money and GDP was identified.

Long-run neutrality of money in the United States in the years 1959–2009 was tested by Lee, who applied a nonparametric testing procedure based on spectral approaches for analyzing relation between nominal money and real output (Lee 2012). In his approach long-run effects between bivariate integrated series were represented as the spectral density matrix of their first-differences evaluated at the zero frequency. The long-run neutrality was reduced to zero power of the cross spectral density function near the origin. The applied procedure enabled to reject the long-run neutrality for M2.

Quite similar results were obtained by Li *et al.*, who performed some empirical tests for Granger causality between U.S. money and output and between the return and volume of the CSI 300 Index in the years 1959–2014 (Lee *et al.* 2016). They applied the generalized cross-spectral distribution function, which enabled to prove that the test statistic captured the nonlinear Granger causality. On the other hand, Albuquerque *et al.* used a Bayesian approach to study the Granger causality between money and GDP growth in the US (Albuquerque *et al.* 2016). The Great Moderation was characterized by strong causality from money to economic growth in the period 1960–2005. After 2005 money had no influence on output growth.

In the case of United States and the debate concerning the concept of money neutrality interesting results were obtained (Sargent, Surico 2011) who showed the instability of low-frequency regression coefficients that Lucas (1980) used to express the quantity theory of money. The authors extended the Lucas analysis (1955–1975) to a long period of 1900–2005. In their research a DSGE model estimated over a subsample like Lucas's implied values of the regression coefficients that confirm Lucas's results. However, perturbing monetary policy rule parameters away from the values estimated over Lucas's subsample alters the regression coefficients in ways that reproduce their instability over longer sample. Thus, the authors were able to document periods in which the quantity theory of money propositions broke down and identified alterations in monetary policies that can account for those breakdowns.

Moving to international experience Bae *et al.* used a bivariate, fractionally integrated, autoregressive, moving average (ARFIMA) model of money and real output to extend Fisher and Seater's long-run neutrality and superneutrality in an ARIMA framework (Bae *et al.* 2005). The research was based on time series for a century of low frequency income and money supply data for Argentina (1884–2001), Canada (1880–2001), Italy (1880–1998), Sweden (1880–1995), the United Kingdom (1880–2001), and the United States (1880–2001). The authors confirmed long-run neutrality in every country except Sweden. On the other hand, form the policy perspective, they found that even when money has no lasting long-run impact on real output, positive monetary shocks tended to have a significant and persistent positive effect on the level of output.

Chen investigated the long-run money neutrality with application of data for South Korea (1970–2004) and Taiwan (1965–2004), where a special attention was given to the integration and cointegration properties of the variables (Chen 2007). The research supported the long-run neutrality of money with respect to real output for South Korea. However, in the case of Taiwan, there was no evidence that could support the long-run monetary neutrality for Taiwan. Additionally, based on his estimations the short-run neutrality of money was rejected for both economies.

An interesting comparative study for developed and developing economies was proposed by Issaoui *et al.,* who applied a structural vector error correction (SVEC) model to a comparative research on the United States, Morocco and Gabon for the years 1960–2011 in the case of the US and Morocco, and 1962–2011 for Gabon (Issaoui *et al.* 2015). They examined the long-term relationship among money supply (M2), income (GDP), and prices (CPI). They showed noticeable differences between developed and developing economies, as they confirmed the long-term non-neutrality of money supply in the United States, and its neutrality in Gabon and Morocco. Their results showed that the short-term effect of a monetary shock on GDP is positive and decreasing, which can be treated as a confirmation of the Keynesian approach, according to which money positively affects output. In the case of the long-term, the effects of the shocks were not systematically zero. They showed that it was zero in Gabon while positive, decreasing, and at lower rates in Morocco and the USA.

Concentrating strictly on the developing countries Wang *et al.* analyzed the time-varying causal dynamics in the case of China's money and output based on Markov switching causality approach (Wang *et al.* 2014). The advantage of this approach is that the Markov switching causality method might capture the time-varying causality patterns in an endogenous way. The results indicated that there was a bidirectional time-varying Granger causality between money and output in China.

Saatcioglu and Korap investigated validity of the quantity theory of money relationship for Turkish economy in the years 1950–2006 with application of some contemporaneous time series estimation techniques (Saatcioglu, Korap 2009). They found that there exists an about one-to-one proportionality between money and prices and money and real income, and that erogeneity of money cannot be rejected for the currency in circulation in the economy. However, in the case of the broad monetary aggregate money seemed to be endogenous as for the long-term variable space.

Asongu analyzed the long-and short-run effects of monetary policy on economic activity (output and prices) in the CEMAC and UEMOA CFA franc zones with application of VARs within the frameworks of Vector Error-Correction Models and Granger causality models (Asongu 2016). Based on the research a hypothesis that monetary policy variables affect prices in the long-run but not in the short-run in the CFA zones was rejected. In regard to the influence of monetary variables on output he confirmed the neutrality of money in the long term. On the other hand, in the short-run with the exception of overall money supply, the significant effect of money on output was more relevant in the UEMOA zone, than in the CEMAC zone, where only financial system efficiency and financial activity were significant.

Telatar and Cavusoglu investigated the hypothesis of long-run neutrality and long-run superneutrality of money in the second half of XX century for high inflation countries: Argentina, Brazil, Ecuador, Mexico, Uruguay and Turkey (Telatar, Cavusoglu 2005). They applied Fisher and Seater's bivariate ARIMA representation where money and real output are modeled with log-linear system (Fisher, Seater 1993). Their final results were country specific as money was long-run neutral but not superneutral with respect to real output for Argentina and Uruguay. In these countries money growth had a negative effect on real output. The long-run superneutrality was confirmed for Brazil, Mexico and Turkey. On the other hand, the long-run neutrality was rejected for Ecuador.

Moreover, Hiscock and Handa concentrated on testing long-run money neutrality and superneutrality for all South American economies for the years 1960–2009 (see Hiscock, Handa 2013). By analogy to previous research they applied Fisher and Seater's procedure and they used M1 and M2 as monetary variables. In their research money neutrality was not rejected for Brazil, Chile, Colombia, Guyana, Suriname, Uruguay and Venezuela. However, it was rejected for Argentina, Bolivia, Ecuador, Paraguay and Peru. In the case of the countries, where superneutrality could be tested, it was not rejected for Bolivia, Brazil, Chile, Colombia, Guyana, and Uruguay. It was rejected for Argentina and Peru.

For Czech Republic, a vector autoregression approach was implemented to support the relationships between output and few macroeconomic variables. The causality approach suggested that past price evolution is a Granger cause for interest rate, previous evolution of GDP influences the interest rate while past real GDP causes variations in price level (Urbanovský 2016).

For Romania and Slovak Republic, a Bayesian approach put into evidence the negative impact of internal credit on output growth in the period 1995–2016 (Korauš *et al.* 2017). This negative correlation might be explained by the changes brought by the global economic crisis. For Romania, the empirical findings for the relationship between GDP and monetary variables showed that there was a bidirectional relation between changes in real money demand and output growth in the period 2000–2015, using data with quarterly frequency (Simionescu *et al.* 2017).

The presented review of empirical literature confirms that the research on the relation between monetary variables and output is still in the center of interests of economists. What is more, the provided literature review confirms important differences in the final results depending on the economy and its stage of development, analyzed period and applied methodological approach.

2. Methodology

The stationarity of a time series is an essential characteristic, because it might affect its behaviour. If the data series for *x* and *y* are non-stationary and random processes (integrated series), when modelling the relationship between x and y, the use of a simple OLS could determine a spurious (false) regression. By definition, the stationary of a time series is a statistical property that implies constant mean and variance over time.

If the data series is stationary without differencing it, then the data series is integrated of order 0 (I(0)). If the data set is stationary in the first difference, then it is integrated of order 1 (I(1)). Augmented Dickey-Fuller or Phillips-Perron test might be used to test for the stationary character of the data series.

According to Simionescu, the number of cointegrated vectors could be determined using Johansen and Juselius cointegration test: trace test or maximum eigenvalue test (Simonescu 2014). The last test considers a null hypothesis that states that there are r cointegrating relations, while the alternative assumption considers r+1 cointegrating relationships when r = 0,1,2,…, n–1. The test statistics has the following form:

$$LR_{\max}\left(\frac{r}{n}+1\right)=-T\log(1-\widehat{\lambda}), \tag{1}$$

where λ – maximum eigenvalue, T – sample size, n, r – number of cointegrating relationships, $LR_{\max}$ – maximum likelihood ratio.

Trace statistics test considers a null hypothesis that states that there are r cointegrating relations, while the alternative assumption considers n cointegrating relationships when r = 0, 1, 2, …, n–1. The test statistics has the following form:

$$LR_{tr}\left(\frac{r}{n}\right)=-T\sum_{i=r+1}^{n}\log\left(1-\widehat{\lambda_i}\right), \tag{2}$$

where λ – maximum eigenvalue, T – sample size, n, r – cointegrating relationships, $LR_{\max}$ – trace likelihood ratio.

In practice, there are cases when maximum eigenvalue and trace statistics indicate different results, but always the result based on trace test should be preferred (see Fałdziński *et al.* 2016; Pietrzak *et al.* 2017).

If a cointegration relationship has been identified, then there is a long-run equilibrium relationship between the variables. The vector error correction model (VECM) is estimated in order to assess the short-run characteristics of the co-integrated series. In case of no cointegrated relationship, the VECM is not estimated and Granger causality is checked on stationary data series. For the VECM we have the following regression equations:

$$\Delta Y_t=\alpha_1+p_1e_1+\sum_{i=0}^{n}\beta_i\Delta Y_{t-i}+\sum_{i=0}^{n}\delta_i\Delta X_{t-i}+\sum_{i=0}^{n}\gamma_i Z_{t-i}; \tag{3}$$

$$\Delta X_t=\alpha_2+p_2e_2+\sum_{i=0}^{n}\beta'_i\,\Delta Y_{t-i}+\sum_{i=0}^{n}\delta'_i\,\Delta X_{t-i}+\sum_{i=0}^{n}\gamma'_i\,Z_{t-i}, \tag{4}$$

where Y_{t-i} , X_{t-i}, Z_{t-i} – variables with lag i, ΔY_t, ΔX_t – variables Y and X in first difference, α_1, α_2 – intercepts, e_1, e_2 – error terms, p_1, p_2, β_i, β'_i, δ_i, δ'_i, γ_i, γ'_i – coefficients.

The cointegration rank in the VECM framework indicates the number of cointegrating vectors. A significant and negative coefficient for VECM shows that the short-run fluctuations between the dependent variable and the explanatory ones determine a stable long-term relationship between variables.

We can describe the Granger causality between two variables *X* and *Y* as:

$$Y_i = \alpha_0 + \alpha_1 Y_{t-i} + \ldots + \alpha_i Y_{t-i} + \beta_1 X_{t-1} + \ldots + \beta_i X_{t-i} + \mu; \tag{5}$$

$$\mathrm{X_t} = X_i = \alpha'_0 + \alpha'_1 X_{t-i} + \ldots + \alpha'_i X_{t-i} + \beta'_1 Y_{t-1} + \ldots + \beta'_i Y_{t-i} + \mu', \tag{6}$$

where X_t, Y_t – variables *X* and *Y* at time *t*, α_0, α_1,..., α_i, $\alpha'_0, \alpha'_1, \ldots, \alpha'_i, \beta_1, \ldots, \beta_i, \beta'_1, \ldots, \beta'_i$ – coefficients, μ, μ' – errors term following a white noise. The subscript is used for time periods.

The trend in *X* and *Y* is seen as general movements of cointegration that appear between *X* and *Y* and it follows a unit root process. We can apply the test twice: to study if *X* does not Granger cause *Y* and to study if *Y* does not Granger cause *X*. If the first null hypothesis is not rejected and the second one is rejected, then changes in *X* are Granger caused by the modifications in *Y*. If the first null hypothesis is rejected and the second one is not rejected, then changes in *Y* are Granger caused by the modifications in *X*. If both assumptions are rejected, then we have a bidirectional relationship between variables.

3. Models for GDP in Romania and Czech Republic considering monetary explanatory variables

The data of this research are related to the following macroeconomic variables for Czech Republic and Romania in the period 1995:Q1 – 2015:Q4: gross domestic product (GDP) at market prices (Chain linked volumes (2010), million Euro), Euro/RON, respectively Euro/Koruna exchange rate (*ER*), monetary policy interest rate, money demand *M2*, money supply *M3* (defined as liquid liabilities of the banking system), and domestic or internal credit. The Eurostat database is the source for GDP data series, the rest of the variables having data provided by the National Bank of the two states.

After the transformation disturbances in the early 90s, Romania has experienced a long period of economic growth, which was interrupted by two international crisis episodes: the international instability of the year 2001 and then the world economic and financial crisis in the last quarter of 2008. As a result, in the analyzed period the first significant shock in the Romanian economy occurred in the second quarter of 2001 because of the international context: the decrease in oil price and in the EU imports, the famous terrorist attacks in the 11th of September 2001, less industrial production in the euro area. In Romania, the industrial production declined because of the capital market contraction in the first three quarters. However, these external threats did not decrease consumption and GDP. The signals of crisis were observed again in the middle of 2007, when the confidence indicator decreased in Romania and in the entire EU. Starting with the middle of 2008, generalized contractions in output were registered in the EU countries. Only in the last quarter of 2008, Romanian economy was declared to be in the recession period, while the foreign direct investment flow decreased since the beginning of 2009. The minimum value of GDP was registered in the third quarter of 2010. A recovery of labour market and GDP was observed starting with the end of 2010.

Unlike Romania, the Czech Republic transformed relatively quickly to a market economy. The conservative monetary and fiscal policies created a stable macroeconomic environment. However, in 1997 and 1998, the Czech Republic registered a serious economic decline, which was a consequence of a financial crisis after a massive speculative attack on the Czech crown in May 1997 conducted by international investors. After a set of serious structural and institutional reforms that were forced by the recession, since 2003 the output has registered a constant increase. The positive trend was supported by the entry in the EU in 2004, which helped to increase competitive advantage of the Czech goods on the foreign markets.

As the data series have quarterly frequency, the data were seasonally adjusted using the Tramo-Seats method. The correlation matrix is built for the variables with seasonally adjusted data (see Table 1), *GDP_SA*, *CREDIT_SA, M2_SA* represent seasonally adjusted data series for *GDP*, internal credit, money demand.

Table 1. Correlation matrix of the variables for Romania (source: authors' calculations)

	GDP_SA	CREDIT_SA	ER	INTEREST	M2_SA	M3
GDP_SA	1.000000	0.584920	0.822936	–0.928452	0.933749	0.900368
CREDIT_SA	0.584920	1.000000	0.811584	–0.600563	0.588593	0.472047
ER	0.822936	0.811584	1.000000	–0.907489	0.808384	0.706380
INTEREST	–0.928452	–0.600563	–0.907489	1.000000	–0.889402	–0.844829
M2_SA	0.933749	0.588593	0.808384	–0.889402	1.000000	0.970475
M3	0.900368	0.472047	0.706380	–0.844829	0.970475	1.000000

Table 1 presents correlation matrix of the variables for Romania. Based on the values of Pearson's coefficients of correlation, in Romania there is a very strong correlation between the following pairs of variables: *GDP* and monetary policy interest rate, *M2* and *GDP* and *M3* and GDP. The negative connection between *GDP* and interest rate is in line with the economic theory, because a lower interest rate promotes investment that is just a part of the *GDP*. The increase in money supply will generate increase in *GDP*. In general, the credit expansion brings increases in *GDP* and other macroeconomic variables. The exchange rate affects the net export. A depreciation of the national currency will increase net export and, consequently, the *GDP*. In this particular case, the national currency depreciation stimulates the deposits' holders to keep the money in foreign currency. For Czech Republic, the correlation matrix is built in Table 2.

Correlation matrix of the variables for Czech Republic is given in Table 2. In Czech Republic, a very strong correlation was observed between the following variables: *GDP* and *M2, GDP* and *M3*, *GDP* and internal credit, *M3* and internal credit, *M2* and internal credit, *M2* and *M3*. *GDP* is well correlated with exchange rate (Euro-Koruna), but weakly correlated with monetary policy interest rate. Taking into account all these connections, we can state that *GDP* in the Czech Republic might be explained by *M3*, internal credit and Euro/Koruna exchange rate.

Table 2. Correlation matrix of the variables for Czech Republic (source: authors' calculations)

	GDP_SA	M2_SA	M3_SA	ER	INTEREST	CREDIT_SA
GDP_SA	1.000.000	0.908422	0.919134	−0.85315	−0.46237	0.928727
M2_SA	0.908422	1.000.000	0.99939	−0.77222	−0.67428	0.991988
M3_SA	0.919134	0.99939	1.000.000	−0.78173	−0.66297	0.993541
ER	−0.85315	−0.77222	−0.78173	1.000.000	0.2884	−0.83909
INTEREST	−0.46237	−0.67428	−0.66297	0.2884	1.000.000	−0.60934
CREDIT_SA	0.928727	0.991988	0.993541	−0.83909	−0.60934	1.000.000

The econometric model will relate *GDP* to *M3* and internal credit. Before estimating the models, the presence of unit roots in data will be checked using Phillips-Perron test (see Table 3).

For Czech Republic, the data series for the following variables are stationary at 5% level of significance: *GDP* rate, *GDP* in first difference, internal credit in first difference, exchange rate in first difference, interest rate in first difference, *M2* in the second difference, *M3* in the first difference, *M3* rate.

For Romania, *GDP*, *M3* and internal credit data series in first difference are stationary at 5% level of significance. These results indicated that the variables are co-integrated of order 1. Moreover, the rates of *GDP, M2, M3* and internal credit have stationary data series at 5% level of significance.

Table 3. Results of Phillips-Perron test for checking the presence of unit roots
(source: authors' calculations)

Romania			Czech Republic		
Data series	Adjusted *t* statistics	p-value	Data series	Adjusted *t* statistics	p-value
GDP_SA in level	−1.881876	0.6550	GDP_SA in level	−1.607592	0.5027
	0.085280	0.9628		−1.575171	0.0663
	2.682684	0.9981		2.251389	0.9936
GDP_SA in first difference	−6.434980	0.0000	GDP_SA in first difference	−4.925436	0.0010
	−6.389570	0.0000		−4.839528	0.0002
	−5.832015	0.0000		−3.985701	0.0001
GDP rate	−9.968784	0.0000	GDP rate	−5.016107	0.0007
	−16.31503	0.0000		−4.796143	0.0002
	−15.86018	0.0001		−2.031178	0.0066
Internal credit	−2.656990	0.2571	Internal credit	−1.826921	0.6776
	−2.063056	0.2600		−0.544113	0.8737
	−0.796843	0.3679		1.939463	0.9865

End of Table 3

Romania			Czech Republic		
Data series	Adjusted *t* statistics	p-value	Data series	Adjusted *t* statistics	p-value
Internal credit in first difference	–8.879891	0.0000	Internal credit in first difference	–3.9528	0.01551
	–8.927502	0.0000		–2.984946	0.0428
	–8.951045	0.0000		–2.564575	0.01097
Internal credit rate	–9.122590	0.0000	Exchange rate in first difference	–7.056189	0.000
	–9.036750	0.0000		–7.062318	0.000
	–8.859773	0.0000		–7.086831	0.000
M3	–1.737481	0.7257	Interest rate in first difference	–4.71542	0.002
	1.008977	0.9964		–4.766422	0.0003
	2.525825	0.9971		–4.679939	0.000
M3 in first level	–8.646517	0.0000	M2 in second difference	–3.155717	0.005
	–8.356452	0.0000		–3.063938	0.025
	–7.821429	0.0000		–2.067259	0.066
Rate of M3	–5.859633	0.0001	M3 in the first difference	–3.177	0.049
	–5.233011	0.0001		–3.1355	0.023
	–3.641660	0.0006		–3.591416	0.04567
Rate of M2	–7.388108	0.0000	M3 rate	–3.5488	0.0438
	–3.715698	0.0006		–3.30	0.01754
	–2.275543	0.0020		–3.10	0.01071

The Granger causality is checked for these variables with stationary data (see Table 4 for Romania and Table 5 for Czech Republic).

Table 4. Granger causality between economic growth and rate of *M3*, respectively rate of internal credit in Romania (source: authors' calculations)

Hypothesis	F computed	p-value
Rate of internal credit does not cause economic growth in Granger approach	0.25518	0.7754
Rate of M3 does not cause economic growth in Granger approach	7.40730	0.0023
Rate of internal credit does not cause rate of M3 in Granger approach	10.8914	0.0002
Rate of M2 does not cause economic growth in Granger approach	0.67685	0.51125

According to Granger causality test, the rate of *M3* in Romania was a cause for economic growth over the period 1995:Q1 – 2015:Q4, while rate of internal credit was an indirect cause through the *M3*. On the other hand, rate of *M2* was not a cause for economic growth in Granger sense.

Table 5. Granger causality between economic growth and rate of *M3*, respectively rate of internal credit in Czech Republic (source: authors' calculations)

Hypothesis	F computed	p-value
Internal credit variation does not cause economic growth in Granger approach	1.99034	0.14733
Rate of M3 does not cause economic growth in Granger approach	0.81118	0.45009
Internal credit variation does not cause rate of M3 in Granger approach	2.04589	0.13995
M2 in double difference does not cause economic growth in Granger approach	0.66471	0.51900

In Czech Republic, there is not a causal relationship between internal credit variation and economic growth or *M3* and economic growth.

The cointegration tests (trace and max-eigenvalue tests) indicated the existence of one co-integrating equation at 5% level of significance. For variables in level an error correction model (ECM) was built in order to study the relationship between variable on long-run and on short-run. According to most of the lag length criteria, the optimal lag equals 2 for Romania and Czech Republic.

The following representation for the ECM model for Romania was proposed:

$$D(GDP\_SA) = C(1)\cdot(GDP\_SA(-1) - 0.04546668181\cdot M2\_SA(-1) - 0.0002471190338\cdot CREDIT\_SA(-1) - 22649.82465\ ) + C(2)\cdot D(GDP\_SA(-1)) + C(3)\cdot D(GDP\_SA(-2)) + C(4)\cdot D(M3\_SA(-1)) + C(5)\cdot D(M2\_SA(-2)) + C(6)\cdot D(CREDIT\_SA(-1)) + C(7)\cdot D(CREDIT\_SA(-2)) + C(8); \tag{7}$$

$$D(M2\_SA) = C(9)\cdot(GDP\_SA(-1) - 0.04546668181\cdot M2\_SA(-1) - 0.0002471190338\cdot CREDIT\_SA(-1) - 22649.82465) + C(10)\cdot D(GDP\_SA(-1)) + C(11)\cdot D(GDP\_SA(-2)) + C(12)\cdot D(M2\_SA(-1)) + C(13)\cdot D(M2\_SA(-2)) + C(14)\cdot D(CREDIT\_SA(-1)) + C(15)\cdot D(CREDIT\_SA(-2)) + C(16); \tag{8}$$

$$D(CREDIT\_SA) = C(17)\cdot(GDP\_SA(-1) - 0.04546668181\cdot M2\_SA(-1) - 0.0002471190338\cdot CREDIT\_SA(-1) - 22649.82465) + C(18)\cdot D(GDP\_SA(-1)) + C(19)\cdot D(GDP\_SA(-2)) + C(20)\cdot D(M2\_SA(-1)) + C(21)\cdot D(M2\_SA(-2)) + C(22)\cdot D(CREDIT\_SA(-1)) + C(23)\cdot D(CREDIT\_SA(-2)) + C(24); \tag{9}$$

$$D(GDP\_SA) = -0.08111963182\cdot(GDP\_SA(-1) - 0.04546668181\cdot M2\_SA(-1) - 0.0002471190338\cdot CREDIT\_SA(-1) - 22649.82465\ ) + 0.3008934718\cdot D(GDP\_SA(-1)) + 0.09925276527\cdot D(GDP\_SA(-2)) + 0.04515480542\cdot D(M2\_SA(-1)) + 0.02327501229\cdot D(M2\_SA(-2)) - 0.000126470858\cdot D(CREDIT\_SA(-1)) + 0.0001396672449\cdot D(CREDIT\_SA(-2)) - 111.3961102; \tag{10}$$

$$D(M2\_SA) = 1.00134049\cdot(GDP\_SA(-1) - 0.04546668181\cdot M2\_SA(-1) - 0.0002471190338\cdot CREDIT\_SA(-1) - 22649.82465\ ) + 0.6594019921\cdot D(GDP\_SA(-1)) + 0.07200811145\cdot D(GDP\_SA(-2)) + 0.04638389391\cdot D(M2\_SA(-1)) - 0.03947964216\cdot D(M2\_SA(-2)) - 0.00436726442\cdot D(CREDIT\_SA(-1)) - 0.0007755791133\cdot D(CREDIT\_SA(-2)) + 3164.617006; \tag{11}$$

$$D(CREDIT\_SA) = 0.1381872051\cdot(GDP\_SA(-1) - 0.04546668181\cdot M2\_SA(-1) - 0.0002471190338\cdot CREDIT\_SA(-1) - 22649.82465\ ) + 10.57678667\cdot D(GDP\_SA(-1)) - 10.32404898\cdot D(GDP\_SA(-2)) - 0.4527590786\cdot D(M2\_SA(-1)) - 0.9736574256\cdot \mathrm{D(M2\_SA(-2))} - 0.01481089373\cdot \mathrm{D(CREDIT\_SA(-1))} - 0.03537148962\cdot \mathrm{D(CREDIT\_SA(-2))} + 7440.457977, \quad (12)$$

where: $C(1)$, $C(2)$,…,$C(24)$ are coefficients in the regressions, *D(GDP_SA), D(M2_SA), D(CREDIT_SA)* represent seasonally adjusted data series for GDP, money demand and internal credit in first difference.

According to estimations, $C(1)$ has a negative value, fact that shows a long-run causality between GDP, internal credit and M3. In order to establish the existence of the short-run causality, the significance of some of the coefficients is checked ($C(4)$, $C(5)$, $C(6)$ and $C(7)$). The results based on OLS estimation are presented in the Table 6.

Table 6. The significance of the model's coefficients for Romania (source: authors' calculations)

	Coefficient	Standard Error	t-Statistic	Prob.
C(1)	–0.081120	0.038934	–2.083543	0.0384
C(2)	0.300893	0.115849	2.597285	0.0100
C(3)	0.099253	0.118145	0.840090	0.4018
C(4)	0.045155	0.019304	2.339141	0.0202
C(5)	0.023275	0.019950	1.166659	0.2446
C(6)	–0.000126	0.001044	–0.121096	0.9037
C(7)	0.000140	0.001045	0.133683	0.8938
C(8)	–111.3961	100.9450	–1.103533	0.2710
C(9)	1.001340	0.228149	4.388985	0.0000
C(10)	0.659402	0.678871	0.971321	0.3325
C(11)	0.072008	0.692326	0.104009	0.9173
C(12)	0.046384	0.113121	0.410039	0.6822
C(13)	–0.039480	0.116907	–0.337701	0.7359
C(14)	–0.004367	0.006120	–0.713597	0.4762
C(15)	–0.000776	0.006122	–0.126682	0.8993
C(16)	3164.617	591.5330	5.349857	0.0000
C(17)	0.138187	4.397549	0.031424	0.9750
C(18)	10.57679	13.08520	0.808301	0.4198
C(19)	–10.32405	13.34455	–0.773653	0.4400
C(20)	–0.452759	2.180393	–0.207650	0.8357
C(21)	–0.973657	2.253374	–0.432089	0.6661
C(22)	–0.014811	0.117964	–0.125554	0.9002
C(23)	–0.035371	0.118006	–0.299743	0.7647
C(24)	7440.458	11401.76	0.652571	0.5147

Only *C(4)* significantly differs from 0, which indicates that there is also a short-run causality from *M3* to *GDP* in Romania.

All the results regarding the estimation of the ECM model are presented in Appendix 1. The errors are homoscedastic at 5% level of significance, the chi-square statistic being 105.4408 (associated p-value is 0.0569). Moreover, the errors are not serial correlated till a lag equaled to 12 at 5% level of significance.

According to variance decomposition for seasonally adjusted *GDP* in Table 7, in the first period after a shock in economy, the changes in *GDP* are caused only by that shock. After two periods, only 97.01% of the variation in *GDP* is due to changes in *GDP*, 2.98% of the variation being due to changes in money demand *M*2. The influence of *M*2 increases in time, arriving to around 12.57% at the 10^{th} lag.

Table 7. Variance decomposition of gross domestic product in Romania
(source: authors' calculations)

Period	*S.E.*	*GDP_SA*	*M2_SA*	*CREDIT_SA*
1	426.5012	100.0000	0.000000	0.000000
2	698.2645	97.01104	2.983764	0.005201
3	956.0067	93.38813	6.599879	0.011986
4	1182.363	91.33128	8.638524	0.030194
5	1381.869	90.02742	9.928052	0.044524
6	1559.770	89.13294	10.81268	0.054379
7	1719.995	88.49293	11.44548	0.061592
8	1865.788	88.01629	11.91673	0.066980
9	1999.728	87.64864	12.28029	0.071072
10	2123.822	87.35639	12.56937	0.074248

Note: *S.E.* is standard error.

The following representation for the ECM model for Czech Republic was given:

$$D(GDP\_SA) = A(1,1)\cdot(B(1,1)\cdot GDP\_SA(-1) + B(1,2)\cdot M3\_SA(-1) + B(1,3)\cdot CREDIT\_SA(-1) + B(1,4)) + C(1,1)\cdot D(GDP\_SA(-1)) + C(1,2)\cdot D(GDP\_SA(-2)) + C(1,3)\cdot D(M3\_SA(-1)) + C(1,4)\cdot D(M3\_SA(-2)) + C(1,5)\cdot D(CREDIT\_SA(-1)) + C(1,6)\cdot D(CREDIT\_SA(-2)) + C(1,7); \quad (13)$$

$$D(M3\_SA) = A(2,1)\cdot(B(1,1)\cdot GDP\_SA(-1) + B(1,2)\cdot M3\_SA(-1) + B(1,3)\cdot CREDIT\_SA(-1) + B(1,4)) + C(2,1)\cdot D(GDP\_SA(-1)) + C(2,2)\cdot D(GDP\_SA(-2)) + C(2,3)\cdot D(M3\_SA(-1)) + C(2,4)\cdot D(M3\_SA(-2)) + C(2,5)\cdot D(CREDIT\_SA(-1)) + C(2,6)\cdot D(CREDIT\_SA(-2)) + C(2,7); \quad (14)$$

$$D(CREDIT\_SA) = A(3,1)\cdot(B(1,1)\cdot GDP\_SA(-1) + B(1,2)\cdot M3\_SA(-1) + B(1,3)\cdot CREDIT\_SA(-1) + B(1,4)) + C(3,1)\cdot D(GDP\_SA(-1)) + C(3,2)\cdot D(GDP\_SA(-2)) + C(3,3)\cdot D(M3\_SA(-1)) + C(3,4)\cdot D(M3\_SA(-2)) + C(3,5)\cdot D(CREDIT\_SA(-1)) + C(3,6)\cdot D(CREDIT\_SA(-2)) + C(3,7). \quad (15)$$

VAR Model – Substituted Coefficients:
===============================

$$D(GDP\_SA) = 0.06619860718 \cdot (GDP\_SA(-1) + 0.0001106598014 \cdot M3\_SA(-1) - 0.0001446134031 \cdot CREDIT\_SA(-1) - 122.4448315\ ) - 0.002524380149 \cdot D(GDP\_SA(-1)) - 0.08128147219 \cdot D(GDP\_SA(-2)) - 2.269653903\text{e}-05 \cdot D(M3\_SA(-1)) - 3.751938115e-06 \cdot D(M3\_SA(-2)) + 2.794035506e-05 \cdot D(CREDIT\_SA(-1)) - 1.108711242e-05 \cdot D(CREDIT\_SA(-2)) + 1.188112996; \quad (16)$$

$$D(M3\_SA) = -323.6771974 \cdot (GDP\_SA(-1) + 0.0001106598014 \cdot M3\_SA(-1) - 0.0001446134031 \cdot CREDIT\_SA(-1) - 122.4448315) + 8745.767384 \cdot D(GDP\_SA(-1)) + 3721.886157 \cdot D(GDP\_SA(-2)) + 0.31139554 \cdot D(M3\_SA(-1)) + 0.4153618406 \cdot D(M3\_SA(-2)) + 0.09684107879 \cdot D(CREDIT\_SA(-1)) - 0.03362931031 \cdot D(CREDIT\_SA(-2)) + 3177.10184; \quad (17)$$

$$D(CREDIT\_SA) = 271.9401909 \cdot (GDP\_SA(-1) + 0.0001106598014 \cdot M3\_SA(-1) - 0.0001446134031 \cdot CREDIT\_SA(-1) - 122.4448315) + 5022.127607 \cdot D(GDP\_SA(-1)) + 3257.436616 \cdot D(GDP\_SA(-2)) - 0.1164114137 \cdot D(M3\_SA(-1)) + 0.133933403 \cdot D(M3\_SA(-2)) + 0.4701303393 \cdot D(CREDIT\_SA(-1)) + 0.2507196312 \cdot D(CREDIT\_SA(-2)) + 5122.60556. \quad (18)$$

According to estimations for Czech Republic, $C(1)$ has a positive value, fact that does not show a long-run causality between *GDP*, internal credit and *M*3. The coefficients $C(4)$, $C(5)$, $C(6)$ and $C(7)$ are not statistically significant and a short-run causality was not identified.

In the next stage, using the data for both countries, a dynamic panel data model was built to explain the *GDP* using the monetary variables (*M*3 and internal credit) as we can see in Table 8.

Table 8. Dynamic panel model with Arrelano-Bover/Blundell-Bond estimator (source: authors' calculations)

Explanatory variable	Coefficient	Standard error	*Z*	*P*>|*z*|
GDP in the previous period	0.6693275	0.0532114	12.58	0.000
Internal credit	–6.234597	1.479727	–4.21	0.000
*M*3	4.532594	1.071852	4.23	0.000
Constant	828591.5	283018.1	2.93	0.003

Note: *Z* is the computed statistic and *P* is the associated p-value.

The estimations' results (Table 8) showed a positive relationship between *GDP* in the current period and *GDP* in the previous period. An increase in *GDP* in the previous period by one unit will generate an increase in the current *GDP* by 0.669 units. Contrary to expectations, the internal credit was negatively correlated to *GDP*. As Leitão showed that there is not any consensus in literature regarding to support that domestic credit stimulates the economic growth (Leitão 2013). Leitão observed a positive correlation between *GDP* and credit using a dynamic panel data for BRIC countries (Brazil, Russia, China, India) and for the European Union member states during the period 1980–2006 (Leitão 2010). Our empirical results are consistent with the findings of Levine and of Hassan, Sanchez and Yu, Korauš *et al.* who showed a negative impact of domestic credit on *GDP* (Levine 1997; Sanchez, Yu 2011; Korauš *et al.* 2017). In this case, the internal credit discouraged the investment and saving that might support the economic growth of a country. Other authors showed that a system using high taxes discourages the economic growth of that country (Padovano, Galli 2002; Koch *et al.* 2005; Lee, Gordon 2005). The fiscal policy could be used as control measure opportunity to adjust inflation or government spending. As expected, *M3* had a positive influence on *GDP*. At each increase in *M3* by 1 unit, the *GDP* increases, in average, by almost 4.53 units.

However, it should be noted that the period from 2002 to 2016 was characterized in both countries by phases of economic growth and recession. The business cycle phases could affect the relationship between *GDP* and monetary variables.

Conclusions

The relationship between output and money has been the subject of many studies in the last years. The results depend on the type of economy. In this research, we chose two countries that experience a different economic development after the collapse of communist regime. Using the VECM framework, we showed that in Romania and in Czech Republic there was a long-run relationship between *GDP*, internal credit and *M3* and a short-run relationship only from *M3* to output. The rate of *M3* was a cause of economic growth in Romania. However, it was not confirmed for Czech Republic. Alternative methods for studying the relationship money-output could be Bayesian VAR models or complex approaches based on DSGE models.

The novelty of this research is given by the comparative analysis of this relationship for two countries with different economic evolutions in the transition from planned economy to market economy. The obtained interesting results contradict the money neutrality hypothesis in post-transformation Central European economies. We brought evidence against money neutrality based on two different econometric approaches: a time series approach for VEC models and a panel data approach using dynamic models. Contrary to economic theory, the internal credit was negatively correlated to *GDP*. In our case, the internal credit discouraged the investment and saving that might support the economic growth of a country.

In that context as expected, the dynamic panel approach indicated that *M3* had a positive impact on output, but the internal credit had a negative influence, as it could discourage investment and saving.

As a result, the provided research can be treated as a voice against the hypothesis of money neutrality in post-transformation countries. The practical implications of these results are related to the policies design. The fiscal policy could be used as control measure opportunity to adjust inflation or government spending.

In the future, this research might be extended in regard to methodological perspective, for example by introducing DSGE models. The additional direction of future research can be seen in the need for comparisons of the results with those for other countries in the Central and Eastern Europe.

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Appendix 1

Table A1. Lag length criteria in the case of the model for Romania (source: authors' calculations)

Lag	LogL	LR	FPE	AIC	SC	HQ
0	–2655.628	NA	1.96E+26	69.05527	69.14659	69.09180
1	–2215.812	833.9363	2.71E+21	57.86525	58.23052*	58.01136
2	–2201.600	25.84109*	2.37E+21*	57.72986*	58.36908	57.98554*
3	–2199.898	2.960686	2.87E+21	57.91944	58.83261	58.28470
4	–2194.340	9.239892	3.16E+21	58.00883	59.19595	58.48367
5	–2187.274	11.19472	3.35E+21	58.05908	59.52015	58.64349
6	–2180.887	9.622257	3.63E+21	58.12694	59.86197	58.82094
7	–2177.258	5.185088	4.25E+21	58.26643	60.27541	59.07001

* indicates lag order selected by the criterion
LR: sequential modified LR test statistic (each test at 5% level)
FPE: Final prediction error
AIC: Akaike information criterion
SC: Schwarz information criterion
HQ: Hannan–Quinn information criterion

Table A2. Lag length criteria in the case of the model for Czech Republic (source: authors' calculations)

Lag	LogL	LR	FPE	AIC	SC	HQ
0	–1588.791	NA	2.45E+22	60.06760	60.17912	60.11048
1	–1234.682	654.7680	5.42E+16	47.04461	47.49071*	47.21616
2	–1220.293	24.97676	4.44E+16*	46.84126*	47.62194	47.14147*
3	–1214.841	8.846478	5.12E+16	46.97515	48.09041	47.40402
4	–1211.439	5.135090	6.44E+16	47.18639	48.63623	47.74393
5	–1198.522	18.03551*	5.71E+16	47.03857	48.82299	47.72477

* indicates lag order selected by the criterion
LR: sequential modified LR test statistic (each test at 5% level)
FPE: Final prediction error
AIC: Akaike information criterion
SC: Schwarz information criterion
HQ: Hannan–Quinn information criterion

Table A3. Vector error correction estimates (source: authors' calculations)

Sample(adjusted): 1995:4 2015:4
Included observations: 81 after adjusting endpoints
Standard errors in ( ) & t–statistics in [ ]

Cointegrating Eq:	CointEq1		
GDP_SA(–1)	1.000000		

Continued Table A3

M2_SA(–1)	–0.045467		
	(0.00347)		
	[–13.0978]		
CREDIT_SA(–1)	–0.000247		
	(0.00278)		
	[–0.08878]		
C	–22649.82		
Error Correction:	D(GDP_SA)	D(M2_SA)	D(CREDIT_SA)
CointEq1	–0.081120	1.001340	0.138187
	(0.03893)	(0.22815)	(4.39755)
	[–2.08354]	[4.38898]	[0.03142]
D(GDP_SA(–1))	0.300893	0.659402	10.57679
	(0.11585)	(0.67887)	(13.0852)
	[2.59728]	[0.97132]	[0.80830]
D(GDP_SA(–2))	0.099253	0.072008	–10.32405
	(0.11815)	(0.69233)	(13.3445)
	[0.84009]	[0.10401]	[–0.77365]
D(M2_SA(–1))	0.045155	0.046384	–0.452759
	(0.01930)	(0.11312)	(2.18039)
	[2.33914]	[0.41004]	[–0.20765]
D(M2_SA(–2))	0.023275	–0.039480	–0.973657
	(0.01995)	(0.11691)	(2.25337)
	[1.16666]	[–0.33770]	[–0.43209]
D(CREDIT_SA(–1))	–0.000126	–0.004367	–0.014811
	(0.00104)	(0.00612)	(0.11796)
	[–0.12110]	[–0.71360]	[–0.12555]
D(CREDIT_SA(–2))	0.000140	–0.000776	–0.035371
	(0.00104)	(0.00612)	(0.11801)
	[0.13368]	[–0.12668]	[–0.29974]
C	–111.3961	3164.617	7440.458
	(100.945)	(591.533)	(11401.8)
	[–1.10353]	[5.34986]	[0.65257]
R–squared	0.198846	0.430752	0.021436
Adj. R–squared	0.122023	0.376167	–0.072399

End of Table A3

Sum sq. resids	13278936	4.56E+08	1.69E+11
S.E. equation	426.5012	2499.277	48173.41
F–statistic	2.588361	7.891344	0.228442
Log likelihood	–601.2273	–744.4467	–984.1100
Akaike AIC	15.04265	18.57893	24.49654
Schwarz SC	15.27914	18.81542	24.73303
Mean dependent	179.2762	3303.807	2787.107
S.D. dependent	455.1752	3164.319	46518.87
Determinant Residual Covariance		2.47E+21	
Log Likelihood		–2327.202	
Log Likelihood (d.f. adjusted)		–2339.837	
Akaike Information Criteria		58.44043	
Schwarz Criteria		59.23857	

Table A4. VEC Residual serial correlation LM tests (source: authors' calculations)

H0: no serial correlation at lag order h
Sample: 1995:1 2015:4
Included observations: 81

Lags	LM–Stat	Prob
1	14.78915	0.0969
2	7.454498	0.5899
3	7.084988	0.6283
4	7.713913	0.5632
5	6.419858	0.6973
6	3.354944	0.9485
7	6.797756	0.6582
8	6.233630	0.7163
9	2.002111	0.9914
10	7.292665	0.6067
11	4.998815	0.8344
12	10.73309	0.2944
Probs from chi–square with 9 df.		

Table A5. VEC residual heteroskedasticity tests (source: authors' calculations)

Sample: 1995:1 2015:4

Included observations: 81

Joint test:			
Chi–sq	df	Prob.	
105.4408	84	0.0569	

Table A6. Unrestricted cointegration rank tests (source: authors' computations)

Hypothesized		Trace	5 Percent	1 Percent
No. of CE(s)	Eigenvalue	Statistic	Critical Value	Critical Value
None *	0.275770	32.94760	29.68	35.65
At most 1	0.080147	6.813286	15.41	20.04
At most 2	0.000573	0.046443	3.76	6.65
*(**) denotes rejection of the hypothesis at the 5%(1%) level Trace test indicates 1 cointegrating equation(s) at the 5% level Trace test indicates no cointegration at the 1% level				
Hypothesized		Max–Eigen	5 Percent	1 Percent
No. of CE(s)	Eigenvalue	Statistic	Critical Value	Critical Value
None **	0.275770	26.13432	20.97	25.52
At most 1	0.080147	6.766842	14.07	18.63
At most 2	0.000573	0.046443	3.76	6.65
*(**) denotes rejection of the hypothesis at the 5%(1%) level Max–eigenvalue test indicates 1 cointegrating equation(s) at both 5% and 1% levels				