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TRANSMISSION EFFECTS IN THE PRESENCE OF STRUCTURAL BREAKS: EVIDENCE FROM SOUTH-EASTERN EUROPEAN COUNTRIES

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Abstract

In this paper, we investigate the monetary transmission mechanism through interest rate and real effective exchange rate channels, for five South-Eastern European countries, namely Bulgaria, Croatia, Greece, Romania and Turkey. Recent unit root and cointegration techniques in the presence of structural breaks in the data are used in the analysis. The empirical results validate the existence of a valid long-run relationship, with parameter constancy, for each of the five sample countries. Additionally, the estimated impulse response functions regarding the monetary variables and the real effective exchange rate converge and follow a reasonable pattern in all cases.

JEL Classification: E43, F15, F42

Keywords: Monetary Transmission Mechanism, Structural Breaks, LM Unit Root Tests, Cointegration Tests, Impulse Responses.

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1. Introduction

The process of integration of the South-Eastern European economies into the European Union (EU) is continuously evolving and has intensified during the last decade. Some of the South-Eastern European countries are either already members of the European Union (EU) or the Eurozone, or associated with the EU; some others are set to become EU members. This implies that developments in the EU affect the above countries in a more systematic way. At the same time, economic transactions in this region have become more significant and systematic, leading banks, enterprises and individuals to extend their activities in the whole region. Thus, there is a need for systematic and detailed research about the economic policies of the countries in this region, especially at this time with the financial and debt crises in the Eurozone. On the one hand, Greece, a Eurozone member since 2001, is in a deep recession with high sovereign debt, and having agreed to two Economic Adjustment Programmes with the ECB-EU-IMF, is fiscally consolidating and faces high unemployment. On the other hand, the emerging economies of the South-Eastern Europe are characterised by relatively high current account deficits and are more vulnerable to a deterioration in the international economy, since they have been negatively affected by the reduction of external demand and the increase in the cost of borrowing from abroad.

In the present paper we attempt to investigate the monetary transmission mechanism for five countries of South-Eastern Europe, namely Bulgaria, Croatia, Greece, Romania and Turkey. For the transition economies (Bulgaria, Croatia and Romania), this investigation is especially important, since it allows us to understand how fast, and to what extent, a change in the central bank's policy instruments influences domestic variables such as inflation. Note that an increasing number of transition economies are already making use of an inflation targeting regime, or are planning to do so. Additionally, it is important to evaluate whether the monetary transmission mechanism operates differently in the transition economies. Coricelli, Égert and MacDonald, (2006) analysed monetary policy transmission in Central and Eastern Europe through four channels: (i) the interest rate channel; (ii) the exchange rate channel; (iii) the asset price channel; and (iv) the broad lending channel. In the present analysis, we focus on the interest rate and real effective exchange rate channels.

The literature on the monetary policy transmission mechanism is quite large and extensive, with both theoretical and empirical papers. Regarding the interest rate channel, there are three approaches. The 'cost of funds' approach examines how market interest rates are transmitted to retail bank interest rates of comparable maturity (De Bondt, 2002); the 'monetary policy' approach directly tests the impact of changes in the policy rate on retail rates (Sander and Kleimeier, 2004a); and a unifying approach that involves two stages, namely the pass-through from the policy rate to market rates and the transmission from market rates to retail rates. Note that interest rate pass-through is usually investigated using an error correction model (ECM) framework. During the last two decades, several researchers have focused on the transition countries of the Central and Eastern Europe. They have largely focused on the asymmetry of the adjustment process, in relation to the Eurozone countries, and the long-run pass through. Regarding the former, their results are mixed (Opiela, 1999; Crespo-Cuaresma, Égert and Reininger, 2004; Horváth, Krekó and Naszódi, 2004; Sander and Kleimeier, 2004b; Égert, Crespo-Cuaresma and Reininger, 2006); regarding the latter their results indicate that both the contemporaneous and long-run pass-through increase over time, while the mean adjustment lag to full pass-through decreases, as more recent data can be used (Crespo-Cuaresma, Égert and Reininger, 2004; Horváth, Krekó and Naszódi, 2004; Sander and Kleimeier, 2004b). The exchange-rate pass-through in the transition economies has also been studied by several researchers, using mainly vector autoregressive (VAR) and vector error-correction (VECM) models (see, for instance, Darvas, 2001; Mihaljek and Klau, 2001; Coricelli, Jazbec and Masten, 2003; Dabušinskas, 2003; Gueorguiev, 2003; Bitâns, 2004; Kara et al., 2005; Korhonen and Wachtel, 2005).

The novelty of this paper lies in the following. Firstly, we use the most recent data from the mid-1990s to 2011, in order to establish a valid long-run relationship for each sample country and to estimate impulse response functions. Secondly, recently developed Lagrange Multiplier (LM) unit root (Lee and Strazicich, 2003) and cointegration tests (Johansen, Mosconi and Nielsen, 2000 and Lütkepohl and Saikkonen, 2000, and their extensions in several recent papers noted below) have been implemented in the analysis. These tests allow for structural breaks in the data. Such breaks are important in this context, since the economic policies implemented in the sample countries are likely to have caused structural shifts in both the levels and trends of particular variables. Additionally, the countries examined are heterogeneous and at different stages of the process of integration into the EU: Bulgaria and Romania joined the EU in 2007 after a long transition period from centrally-planned to free market economies; Croatia joined the EU in 2013 having also followed a long transition period; Greece has been a Eurozone member since 2001; and Turkey agreed to a customs union with the EU in 1996, is under negotiations for future EU membership, and has also had a stand-by agreement with the IMF for a number of years.

In summary, the empirical evidence validates the existence of structural breaks and identifies a valid long-run relationship among industrial production, the consumer price index, the money supply, the money market rate and the real effective exchange rate, for each of the five countries under consideration. Additionally, the estimated impulse response functions for the monetary variables and the real effective exchange rate converge and seem reasonable in all cases.

The rest of the paper is organised as follows. Section 2 describes briefly the theoretical framework of the analysis and outlines the unit root and cointegration tests in the presence of structural breaks. Section 3 describes the data and analyses the empirical results, while Section 4 provides some concluding remarks.

2. Theoretical framework

In the present study, we estimate a reduced-form model in order to investigate the monetary transmission mechanism for Bulgaria, Croatia, Greece, Romania and Turkey. The analysis will focus on the interest rate channel and the real effective exchange rate channel. We do not attempt to construct a full structural model in order to capture relationships proposed by economic theory, due to (a) data limitations, and (b) the extreme heterogeneity of the sample countries. Thus, our analysis will be based on unit root and cointegration testing in the presence of structural breaks, along with VECM specification and impulse response estimation. Note that structural breaks are important in this context, since the economic policies implemented in the sample countries are likely to have caused structural shifts in the level and trend of the variables under consideration.

2.1 Unit root tests with structural breaks

In order to test the statistical properties of the data, we used the two-break LM (Lagrange Multiplier) test developed by Lee and Strazicich (2003). This test has several desirable properties: (a) it determines the structural breaks "endogenously" from the data; (b) its null distribution is invariant to level shifts in a variable; and (c) it is easy to interpret. By including breaks under both the null and alternative hypotheses, a rejection of the null hypothesis of a unit root implies unambiguously trend stationarity of the variable concerned.

Consider for instance the two-break LM unit root test for the process y_t generated by

$$y_t = \delta' Z_t + e_t, \qquad e_t = \beta e_{t-1} + A(L)\varepsilon_t, \qquad \varepsilon_t \sim iid(0, \sigma^2) \tag{1}$$

where *A(L)* is a *k*-order polynomial and Z_t is a vector of exogenous variables, whose components are determined by the type of breaks in y_t. Lee and Strazicich (2003) extend Perron's (1989, 1993) single-break models to include two breaks in the level (Model A) and two breaks in both the level and trend (Model C) of y_t. Eq. (1) shows that y_t has a unit root if $\beta = 1$, while it is trend stationary if $\beta < 1$. According to the LM principle, a unit root test statistic can be obtained from the test regression if:

$$\Delta y_t = \delta' \Delta Z_t + \phi \tilde{S}_{t-1} + \sum_{i=1}^{k} \theta_i \Delta \tilde{S}_{t-i} + u_t , \tag{2}$$

where $\tilde{S}_t = y_t - \tilde{\psi}_x - Z_t \tilde{\delta},\ t = 2, ..., T$, in which $\tilde{\delta}$ is a vector of coefficients in the regression of Δy_t on ΔZ_t and $\tilde{\psi}_x = y_1 - Z_1 \tilde{\delta}$, where y_1 and Z_1 are the first observations of y_t and Z_t, respectively, and u_t is a white noise error term. The lagged differences of $\tilde{S}_{t-i}$ correct for serial correlation in u_t. The unit root null hypothesis is described by $\phi = 0$ in eq. (2) and is tested by the LM test statistic:

$$\tilde{\tau} = t\text{-statistic for the hypothesis } \phi = 0. \tag{3}$$

To endogenously determine the location of the two breaks $(\lambda_j = T_{Bj}/T,\ j = 1,2)$ the two-break minimum LM test statistic is determined by a grid search over λ:

$$LM_\tau = \inf_\lambda \{\tilde{\tau}(\lambda)\} \tag{4}$$

The critical values for this test are invariant to the break locations (λ_j) for Model A but depend on the break locations for Model C.

2.2 Cointegration tests with structural breaks

As in the case with unit root testing, structural breaks in the data can distort substantially standard inference procedures for cointegration. Thus, it is necessary to account for possible breaks in the data before inference on cointegration can be made. In the recent literature on cointegration in a VAR framework, there are two main approaches that test for cointegration in the presence of structural breaks.

The first approach has been developed by Johansen, Mosconi and Nielsen (2000) (JMN). It extends the standard VECM with a number of additional dummy variables in order to account for q possible exogenous breaks in the levels and trends of the deterministic components of a vector-valued stochastic process. JMN then derive the asymptotic distribution of the likelihood ratio (LR) or trace statistic for cointegration and obtain critical values or p-values, for the multivariate counterparts of models A and C above with q possible breaks, using the response surface method.

To illustrate the JMN approach, consider briefly the simple case with only level shifts in the constant term μ of an observed $p-$dimensional time series y_t, $t = 1,...,T$, of possibly $I(1)$ variables. JMN divide the sample observations into q sub-samples, according to the location of the break points, and assume the following VECM(k) for y_t conditional on the first k observations of each sub-sample $y_{T_{j-1}+1},...,y_{T_{j-1}+k}$:

$$\Delta y_t = \Pi y_{t-1} + \mu D_t + \sum\nolimits_{i=1}^{k-1} \Gamma_i \Delta y_{t-i} + \sum\nolimits_{i=1}^{k} \sum\nolimits_{j=2}^{q} g_{ji} D_{j,t-i} + \varepsilon_t, \quad \varepsilon_t \sim iidN(0,\Omega), \tag{5}$$

where $\mu = (\mu_{1,.........,}\mu_q)$ and $D_t = (D_{1,t,.........,}D_{q,t})^{/}$ are of dimension $(p \times q)$ and $(q \times 1)$, respectively, and the $D_{j,t}$'s are dummy variables, such that $D_{j,t} = 1$ for

$T_{j-1}+k+1 \leq t \leq T_j$ and $D_{j,t}=0$ otherwise, for $j=1,....,q$. The hypothesis of at most r_0 cointegrating relations $\left(0 \leq r_0 < p\right)$ among the components of y_t can be stated in terms of the reduced rank of the $(p \times p)$ matrix $\Pi = \alpha\beta^{/}$, where α and β are matrices of dimension $(p \times r)$. The cointegration hypothesis can then be tested by the likelihood ratio statistic:

$$LR_{JMN} = -T\sum_{i=r_0+1}^{p} \ln\left(1-\hat{\lambda}_i\right) \tag{6}$$

where the eigenvalues $\hat{\lambda}_j$'s can be obtained by solving the related generalized eigenvalue problem, based on estimation of the VECM(*k*) in equation (5), under the additional restrictions that $\mu_j = \alpha\rho_j^{/}$, $j=1,.....,q$, where ρ_j is of dimension $1 \times r$. These restrictions are required in order to eliminate a linear trend in the level of the process y_t (Johansen *et al.*, 2000, p. 218).

The second approach has been developed by Lütkepohl and his associates (Lütkepohl and Saikkonen, 2000; Saikkonen and Lütkepohl, 2000; Trenkler, Saikkonen and Lütkepohl, 2008) (henceforth the LST approach). These authors assume that the DGP for a vector-valued process y_t is such that its deterministic part does not affect its stochastic part. It is then possible to remove the deterministic part, with possible breaks, in the first stage, and carry out Likelihood Ratio (LR) or Lagrange Multiplier (LM) cointegration tests in the second stage using the *de-trended* stochastic part of y_t.

Briefly, in the LST approach the DGP for y_t is the sum of a deterministic part μ_t and a stochastic part x_t, where x_t is an unobservable zero-mean purely stochastic VAR process. Structural shifts in y_t are accounted for by the use of appropriate dummy variables in the deterministic component μ_t. To illustrate the LST approach for LR-type tests, consider the case of a single shift in both the level and the trend of y_t, at time T_B. LST specify the following DGP for y_t:

$$y_t = \mu_t + x_t = \mu_0 + \mu_1 t + \delta_0 d_t + \delta_1 b_t + x_t, \quad t=1,....,T, \tag{7a}$$

where t is a linear time trend, μ_i $(i=0,1)$ and δ_i $(i=0,1)$ are unknown $(p \times 1)$ parameter vectors, d_t and b_t are dummy variables defined as $d_t = b_t = 0$ for $t < T_B$, and $d_t = 1$ and $b_t = t - T_B + 1$ for $t \geq T_B$. The unobserved stochastic error x_t is assumed to follow a $VAR(k)$ process with VECM representation:

$$\Delta x_t = \Pi x_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta x_{t-i} + \varepsilon_t \quad , \quad \varepsilon_t \sim iidN(0,\Omega), \quad t = 1,\ldots,T \ . \tag{7b}$$

It is also assumed that the components of x_t are at most integrated of order one processes and cointegrated (i.e. $\Pi = \alpha\beta^{/}$) with cointegrating rank r_0.

Given the DGP in (7a) and (7b), the first step of the LST approach involves obtaining estimates of the parameter vectors μ_0, μ_1, δ_0 and δ_1 in (7a) using a feasible GLS procedure under the null hypothesis $H_0(r_0): rank(\Pi) = r_0$: vs. $H_1(r_0): rank(\Pi) > r_0$ (see Saikkonen and Lütkepohl (2000) for details). Having the estimated parameters, $\hat{\mu}_0$, $\hat{\mu}_1$, $\hat{\delta}_0$ and $\hat{\delta}_1$, one then computes the de-trended series $\hat{x}_t = y_t - \hat{\mu}_0 - \hat{\mu}_1 t - \hat{\delta}_0 d_t - \hat{\delta}_1 b_t$. In the second step, an LR-type test for the null hypothesis of cointegration is applied to the de-trended series. This involves replacing x_t by $\hat{x}_t$ in the VECM (7b) and computing the LR or trace statistic:

$$LR_{LST} = -T \sum_{i=r_0+1}^{p} \ln(1 - \tilde{\lambda}_i) \ , \tag{8}$$

where the eigenvalues $\tilde{\lambda}_i$'s can be obtained by solving a generalized eigenvalue problem, along the lines of Johansen (1988).

Under the null hypothesis of cointegration, Trenkler *et al.* (2008) derive asymptotic results and *p*-values for the case of one level shift and one trend break in the y_t process, and show that, in this case, the asymptotic distribution of the LR statistic in (8) depends on the location of the break point. They also discuss how the results can be extended to the general case of $q > 1$ break points. Also, critical or $p-$values for a single level shift can be computed by the response surface techniques developed in Trenkler (2008).

Since the JMN and LST approaches have different finite sample properties, we employ both the LR_{LST} and LR_{JMN} test statistics in the subsequent analysis. It is worth noting here that Lütkepohl, Saikkonen and Trenkler (2003) studied the statistical properties of their tests in the case of shifts in the level of y_t and compare them to alternative tests developed by Johansen *et al.* (2000). They found that their tests have better size and power properties than the Johansen et al. tests in finite samples. For that reason, if the results of the JMN and LST tests are different, we will use those of the latter test. The break points are determined from the data on the basis of the results of the two-break LM unit root test discussed above.

3. Data and empirical results

3.1 Data

Our sample consists of monthly data ending 2011:07. The starting date of the data for each country is different, depending on data availability. The time span for Bulgaria begins in 2000:01, for Croatia and Romania in 2002:01, for Greece in 1995:01, and for Turkey in 2003:01. We use data for industrial production (IP), the consumer price index (CPI), the money supply (M1 for Croatia, M2 for Romania, M3 for Bulgaria and Turkey, while for Greece, which is a Eurozone member we do not use the money supply in the analysis), the money market rate (MMR) for all countries except Greece, for which we used the Treasury bill rate (TB), and real effective exchange rates based on consumer prices (REER). All data were obtained from the International Financial Statistics of the IMF, except for the real effective exchange rate for Turkey that was obtained from the Central Bank of Turkey. All data, except interest rates, are transformed into natural logarithms.

3.2 Unit root tests results

Before proceeding to our analysis, each time series is first tested for a unit root. Table 1 reports the unit root results from the two-break LM test. Each time series was tested for a unit root using the two-break LM test at the 1- and 5 percent levels of significance. The number of lags, k, in equation (2) was determined using a “general to specific" procedure at each combination of relative break points (λ_1, λ_2). Initially, the lag-length was set at $k = 12$, and the significance of the last lagged term was examined at the 10 percent level. The procedure was repeated until the last lagged term was found to be significantly different than zero, where the procedure stops.[1]

As shown in the last column of table 1, the unit root hypothesis with two structural breaks cannot be rejected for all variables under consideration. Column 5 of table 1, which presents the estimated structural breaks in each time series, indicates that the consumer price index, the money supply and the Croatian money market rate experience one structural break. Column 3 of table 1 also reports that Model C (i.e. break(s) in both the level and the trend) fits the data best for all cases, over the sample period. Not surprisingly, the estimated structural breaks correspond well to specific events that have taken place in the sample countries during the sample period.

More specifically, the industrial production, the money market rate and the real effective exchange rate of Bulgaria experience a structural break in the 2008-2010 period, which is probably related to the consequences of the global financial crisis. The real effective exchange rate of the country, along with the consumer price index and the money supply appear to have a break in 2007, when Bulgaria became a full member of the EU. Also, the industrial production, the consumer price index, the money supply and the money market rate of Bulgaria experience a structural break in the 2001-2005 period. In general, these breaks can be attributed to certain measures that the country adopted during the long transition period and the negotiations for EU accession. Note that following the 1997 economic and financial crisis, Bulgaria adopted a euro-based currency board to stabilise its exchange rate, and implemented a comprehensive economic plan, which included trade and price liberalisation, welfare sector reform, and divesting state-owned enterprises.

For Croatia, all variables experience a structural break during the 2008-2010 period, again attributable to the consequences of the global financial crisis. Industrial production and the real effective exchange rate of the country appear to have a second

[1] The two-break and one-break LM tests were computed using the Gauss codes of J. Lee available at the website http://www.cba.ua.edu/~jlee/gauss .

break in 2006-2007. During that period, the EU-Croatia negotiations for full membership were started and the process of screening 35 acquis chapters was completed.

The global financial crisis had, of course, a significant impact on the Greek economy. The structural breaks in industrial production and the Treasury bill rate in mid-2008 confirm the above argument. Greek industrial production, the consumer price index and the real effective exchange rate experience a break in early 1999, which is probably related to the formation of the Eurozone. The consumer price index and the real effective exchange rate exhibit a break in the 2001-2002 period, which can be attributed to Greece's membership in the Eurozone and the subsequent adjustments in the country's economy. Finally, the structural break in the Greek Treasury bill rate in early 2004 coincides with an increased budget deficit due to preparations for the Olympic Games and the forthcoming elections.

Moving to Romania, the two structural breaks in industrial production and the second break in the money market rate can be attributed to the global financial crisis. Additionally, the Romanian money supply and real effective exchange rate appear to have breaks in 2007, when the country became a full member of the EU. Both structural breaks in the country's consumer price index, along with the first break in the money supply, the money market rate and the real effective exchange rate occur in the 2003-2006 period. In general, these breaks are a result of measures that the country adopted during its long transition period and the negotiations for EU accession. Note that since 2000, Romania has implemented tight fiscal and monetary policies along with structural reforms designed to support growth and improve financial discipline in the private sector. These reforms have placed the country's public finances and the financial system on a firmer footing. Further, Romania is currently considering a currency board vis-à-vis the euro, in order to reduce inflation and gain monetary policy credibility.

In the case of Turkey, the industrial production and the real effective exchange rate appear to have two breaks during the 2008-2009 period. Note that it is not clear if these breaks are a consequence of the global financial crisis or the negotiations to end the IMF stand-by agreement. After the severe economic crisis that the country faced in 2000-2001, Turkey implemented an IMF programme based on high interest rates in order to attract foreign capital, accompanied by fiscal contraction and privatisations. This programme led to the country's currency ('lira') becoming overvalued and to an import boom in both consumption and investment goods. As a result, Turkey's external indebtedness increased and the deficit on the current account rose to 7.5 per cent of GNP by mid-2008. The structural breaks in the country's consumer price index, money supply and money market rates during 2004-2007 can be attributed to the economic measures adopted in the context of the IMF programme.

3.3 Cointegration tests results

In this section we examine the cointegration results with structural breaks on our reduced-form vector. These results are based on the JMN and the LST procedures described in Section 2.2. As breaks for each country, we use the estimated structural breaks appeared most frequently in table 1. We also avoid using breaks very close to the beginning or the end of our sample. In the case of the JMN procedure we estimate the VECM in equation (5) for each country and computed the LR_{JMN} test statistics and the corresponding response surface p-values using the JMulti software. The Akaike's information criterion is also used to select the appropriate lag length, k, in the VECM for each of the five countries. In the case of the LST procedure, we estimate the model in equations (7a) and (7b) by adjusting (7a) to account for the structural breaks specific to each country. Since all five countries experience two significant breaks in both the level and the trend of their exchange rates, we extend equation (7a) by adding a second step dummy and a second linear trend dummy. Then, for each country we compute the LR_{LST} test statistic and the corresponding response surface p-values using GAUSS routines.[2]

Table 2 reports the LR_{JMN} and LR_{LST} test statistics and the respective p-values, for each of the five sample countries. As shown in the table, the JMN test indicates four cointegrating vectors for Bulgaria, three cointegrating vectors for Croatia, and two for Greece, Romania and Turkey, either at the 5 or at 10 percent level of significance. By

[2] The authors are grateful to Carsten Trenkler for kindly providing them with the Gauss codes for these estimations.

contrast, the LST test indicates a single cointegrating vector in each case.[3] As noted in Section 2.2, the LST test has better size and power properties than the JMN test in finite samples. Thus, our subsequent analysis will be based on the results of the LST test.

Note here, that the JMN and LST tests for cointegration in the presence of structural breaks, assume that the ‘‘long-run’’ cointegration parameters remain constant over the sample period. Otherwise, the test results and inference would be invalid. To test for parameter constancy, we use the methodology developed by Hansen and Johansen (1999), who suggest a graphical procedure based on recursively-estimated eigenvalues. Figure 1 shows the time path of each eigenvalue (i.e. the τ_{sum}-statistic) for the null hypothesis that it is stable. The dotted line in each plot corresponds to 1.36, which is the 5 percent critical value for the Hansen and Johansen parameter constancy test. The null hypothesis of long-run parameter constancy cannot be rejected in all cases, as the time paths of the eigenvalues are always below the dotted line.[4]

3.4 VECMs and orthogonal impulse responses

Based on the cointegration results of the previous section, we have established a valid relationship, which can be interpreted as the long-run relationship between the industrial production, the consumer price index, the money supply, the money market rate and the real effective exchange rate. Following the above, we estimate the corresponding VECMs, based on equations (7a) and (7b). Table 3 presents the estimated coefficients of the solved cointegrating vectors (i.e. reduced form equations) normalised on industrial production, along with the results from the long-run exclusion test. As shown in table 3, most of the estimated coefficients have the expected signs. The long-run exclusion test investigates whether any of the variables can be excluded from the cointegrating space. Using the likelihood ratio test statistic, our results imply that the consumer price index can be excluded from the cointegrating equation for Bulgaria, while both the consumer price index and the real effective exchange rate can be excluded from the cointegrating equation for Croatia. For Greece no variable can be excluded from the cointegration

[3] We have also performed the cointegration tests using diferrent combinations of structural breaks. The estimated results did not change and are available upon request.

[4] We have included centred seasonal dummies in the Hansen-Johansen tests, the VECM estimations and the impulse responses.

space. The consumer price index and the money market rate can be excluded from the cointegrating equation for Romania, while for Turkey, the money supply, the money market rate and the real effective exchange rate can be excluded from the cointegration space. When it comes to the implied structural breaks, the long-run exclusion test shows that none of the breaks can be excluded from the cointegrating space for Croatia and Romania. On the contrary, both structural changes are found statistically insignificant for Bulgaria, Greece and Turkey in the long run.

We also perform weak exogeneity tests, in order to investigate whether a variable can be considered as weakly exogenous to the long-run parameters. A variable is said to be weakly exogenous if the corresponding adjustment coefficient cannot be statistically different from zero. The results for this test are reported in table 4 and provide information on the variables that drive the system to long-run equilibrium. Starting from the case of Bulgaria, the money supply is found to be weakly exogenous and, thus, drives the system to its long-run equilibrium. For Croatia, the consumer price index and the real effective exchange rate are found to be weakly exogenous, while for Greece, the driving forces of the system are industrial production and the real effective exchange rate. For Romania, industrial production and the consumer price index are found to be weakly exogenous, while for Turkey weak exogeneity is established for the consumer price index and the money supply.

Finally, and in order to complete our analysis for the monetary transmission mechanism in each country, we estimate orthogonal impulse response functions, based on a one standard deviation innovation for each of the monetary variables (money supply and money market rate or Treasury bill rate), as well as the real effective exchange rate. The estimated impulse responses, along with their 90% bootstrap confidence bands, are presented in figures 2 to 6.[5] As shown, for most, the range of values is small. In general, they converge in all cases, implying our model is stable, and seem reasonable. Only in the case of Turkey, do the response of money market rate to a shock in money supply and the response of industrial production to a shock in the real effective exchange rate not

[5] The 90% bootstrap confidence bands have been computed by simulations using 1000 replications.

converge to a stable level. A possible explanation for this peculiar result could be the strong inflationary tendencies in the Turkish economy.

4. Concluding remarks

In this study, we investigated the transmission mechanism for five South-Eastern Europe countries, namely Bulgaria, Croatia, Greece, Romania and Turkey. We focused on the monetary transmission through the interest rate and real effective exchange rate channels. Data limitations and the extreme heterogeneity of the above countries did not allow us to construct a full structural model based on economic theory. Thus, we used a small reduced-form model for each country, consisting of five endogenous variables, in order to establish a valid long-run relationship and to analyse the impulse response functions. We also included structural shifts in our analysis, since the economic policies implemented in the sample countries are likely to have caused structural shifts in the level and trend of their variables.

The unit root test indicates the presence of one or two structural breaks for each variable. The cointegration test results in the presence of structural breaks show evidence of a single cointegrating vector with parameter constancy, among industrial production, the consumer price index, the money supply, the money market rate and the real effective exchange rate, for each of the five countries under consideration. These results identify a long-run relationship among the above variables, while the estimated impulse response functions following shocks to the monetary variables and the real effective exchange rate converge and seem reasonable in all cases. The present analysis regarding the monetary transmission mechanism could be extended with the use of a global modelling framework, based on the Global VAR (GVAR) model, which avoids the limitations that arise from the use of single VAR and VECMs models and provides a consistent and flexible framework.

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Table 1: Two-break minimum LM unit root test results

Country	Variable	Model	$\hat{k}$	$\hat{T}_B$	$\hat{\lambda}_1,\ \hat{\lambda}_2$	$LM-$statistic
Bulgaria	IP	C	12	2003:12, 2008:08	0.4, 0.8	-5.4237
	CPI	C	12	2002:04, 2007:10	0.2, 0.6	-4.4316
	M3	C	12	2005:02, 2007:08	0.4, 0.6	-4.4335
	MMR	C	10	2001:12, 2009:02	0.2, 0.8	-4.0249
	REER	C	2	2007:10, 2010:01	0.6, 0.8	-5.0334
Croatia	IP	C	11	2006:02, 2008:10	0.4, 0.8	-5.6393
	CPI	C	12	2006:03^{n}, 2008:01	0.4, 0.6	-4.5633
	M1	C	12	2005:03^{n}, 2008:11	0.4, 0.8	-3.9050
	MMR	C	1	2008:02^{n}, 2008:11	0.6, 0.8	-5.6928
	REER	C	1	2007:11, 2010:01	0.6, 0.8	-5.5588
Greece	IP	C	11	1999:12, 2008:08	0.4, 0.8	-5.5307
	CPI	C	10	1999:02, 2001:10	0.2, 0.4	-5.0288
	TB	C	6	2004:02, 2008:09	0.6, 0.8	-3.9003
	REER	C	12	1999:02, 2002:11	0.2, 0.4	-5.2925
Romania	IP	C	12	2008:08, 2010:03	0.6, 0.8	-5.4496
	CPI	C	10	2003:07, 2005:01	0.2, 0.4	-4.3414
	M2	C	12	2004:12, 2007:11	0.4, 0.6	-4.8818
	MMR	C	6	2006:09, 2009:03	0.4, 0.8	-4.7614
	REER	C	1	2004:11, 2007:10	0.4, 0.6	-4.4244
Turkey	IP	C	12	2008:09, 2009:10	0.6, 0.8	-5.5628
	CPI	C	12	2004:10, 2007:11	0.2, 0.6	-5.2050
	M3	C	6	2005:10, 2007:07	0.4, 0.6	-5.1521
	MMR	C	3	2004:10, 2006:10	0.2, 0.4	-5.2923
	REER	C	1	2008:04, 2009:12	0.6, 0.8	-4.7017

Break Points	Critical values for Model C	
$\lambda=(\lambda_1,\ \lambda_2)$	1%	5%
λ=(0.2, 0.4)	-6.16	-5.59
λ=(0.2, 0.6)	-6.41	-5.74
λ=(0.2, 0.8)	-6.33	-5.71
λ=(0.4, 0.6)	-6.45	-5.67
λ=(0.4, 0.8)	-6.42	-5.65
λ=(0.6, 0.8)	-6.32	-5.73

$\hat{k}$ is the estimated number of to correct for serial correlation. $\hat{T}_B$ denotes the estimated break points. $\hat{\lambda}_1$ and $\hat{\lambda}_2$ are the estimated relative break points. IP stands for industrial production, CPI for consumer price index, M1, M2 and M3 for money supply, MMR for money market rate, TB for Treasury bill rate, and REER for real effective exchange rate. ^{n} indicates no significant break at the 10 percent level of significance. The critical values are from table 2 of Lee and Strazicich (2003).

Table 2. The JMN and LST cointegration tests with structural breaks

Country	$(p-r_0)$	$LR_{JMN}(r_0)$	$LR_{LST}(r_0)$	p-values JMN	p-values LST	$\hat{k}$
Bulgaria	5	233.75**	95.90**	0.000	0.001	9
(breaks on:	4	150.55**	41.77	0.000	0.428	
2005:02, 2007:08)	3	94.57**	17.49	0.001	0.916	
	2	51.26**	7.58	0.023	0.938	
	1	20.88	3.32	0.140	0.756	
Croatia	5	195.25**	93.81**	0.000	0.001	2
(breaks on:	4	118.78**	37.92	0.004	0.624	
2006:02, 2008:11)	3	71.95*	20.42	0.080	0.796	
	2	36.64	9.70	0.379	0.828	
	1	16.77	5.08	0.364	0.503	
Greece	4	164.08**	54.90**	0.000	0.048	8
(breaks on:	3	84.99**	11.92	0.003	0.996	
1999:12, 2008:08)	2	42.12	4.01	0.118	0.998	
	1	13.21	2.48	0.551	0.838	
Romania	5	178.26**	81.54**	0.000	0.017	1
(breaks on:	4	109.44**	40.26	0.025	0.500	
2004:12, 2007:11)	3	69.62	17.58	0.123	0.915	
	2	36.25	3.27	0.414	0.999	
	1	16.75	0.02	0.380	0.999	
Turkey	5	162.06**	73.26*	0.001	0.084	1
(breaks on:	4	99.43*	34.87	0.090	0.765	
2004:10, 2008:08)	3	61.17	27.50	0.309	0.356	
	2	34.08	8.16	0.471	0.912	
	1	12.42	1.07	0.668	0.987	

$\hat{k}$ denotes the estimated lag length in the VECM. ** and * denote rejection of the null hypothesis at the 0.05 and the 0.10 level of significance, respectively.

Table 3. Estimated coefficients of the solved cointegrating vectors

Parameter estimates	Bulgaria	Croatia	Greece	Romania	Turkey
β_{IP}	1.000	1.000	1.000	1.000	1.000
β_{CPI}	2.926	0.162	-7.478**	2.166	-1.697**
	(0.190)	(0.624)	(0.000)	(0.288)	(0.006)
β_{M}	1.439*	0.533**	NA	-2.253**	-0.081
	(0.073)	(0.000)		(0.000)	(0.113)
β_{IR}	-0.137**	-0.004**	-0.027**	-0.016	0.003
	(0.009)	(0.001)	(0.006)	(0.128)	(0.065)
β_{REER}	5.783**	-0.009	-1.553**	4.769**	0.196
	(0.000)	(0.970)	(0.002)	(0.000)	(0.106)
Trend	-0.050**	-0.001	0.024**	0.043**	0.017**
	(0.001)	(0.422)	(0.000)	(0.006)	(0.000)
SB_1	-0.001	-0.001**	0.001	-0.038**	0.003
	(0.787)	(0.002)	(0.720)	(0.000)	(0.153)
SB_2	0.003	-0.003**	-0.001	-0.046**	0.002
	(0.212)	(0.000)	(0.931)	(0.000)	(0.465)

$\beta's$ are the parameters of the solved cointegrating vectors, normalised on the industrial production. M stands for M1, M2 or M3 depending on the country, while IR stands for MMR or TB depending on the country. SB_1 and SB_2 are the first and the second structural trend break, respectively. Numbers in parentheses are the p-values of the likelihood ratio test statistics for the long-run exclusion tests. NA stands for not available. ** (*) denotes rejection of the null hypothesis at the 0.05 (0.10) level of significance.

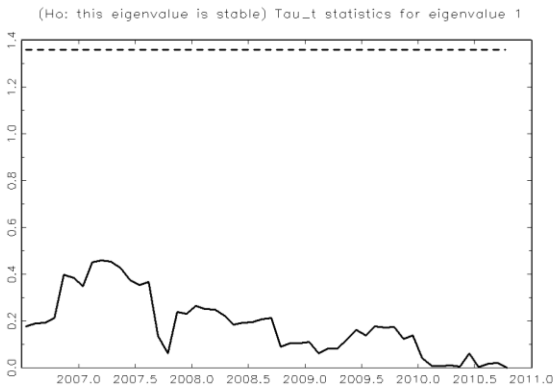
Table 4. Adjustment coefficients and weak exogeneity tests

Parameter estimates	Bulgaria	Croatia	Greece	Romania	Turkey
α_{IP}	-0.067*	-0.642*	0.028	-0.001	-0.455*
	(0.032)	(0.165)	(0.029)	(0.023)	(0.097)
α_{CPI}	0.025*	0.014	-0.015*	0.003	-0.022
	(0.006)	(0.021)	(0.004)	(0.002)	(0.012)
α_{M}	-0.017	0.274*	NA	-0.034*	0.077
	(0.013)	(0.129)		(0.007)	(0.119)
α_{IR}	1.849*	-0.458*	-0.952*	-1.403*	0.581*
	(0.675)	(0.085)	(0.294)	(1.436)	(0.143)
α_{REER}	0.058*	-0.016	-0.006	0.014*	-0.175*
	(0.008)	(0.043)	(0.009)	(0.007)	(0.069)

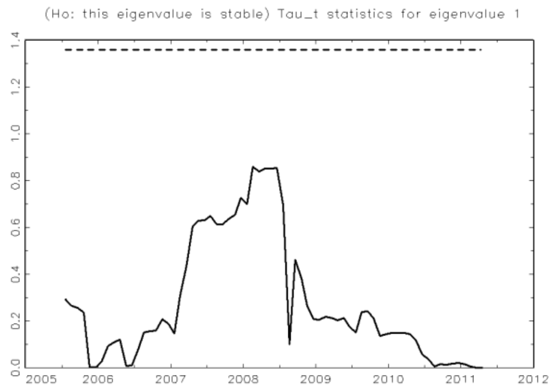
$\alpha's$ are the adjustment coefficients. M stands for M1, M2 or M3 depending on the country, while IR stands for MMR or TB depending on the country. Numbers in parentheses are standard errors. NA stands for not available. * denotes rejection of the null hypothesis $H_0 : \alpha_i = 0$ at the 0.05 level of significance.

Figure 1. Parameter constancy tests (τ_{sum} -statistics)

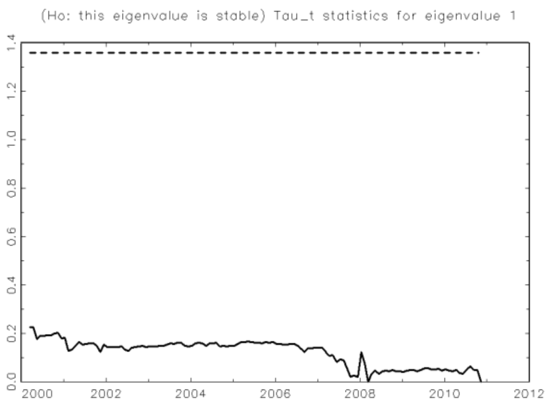
Bulgaria



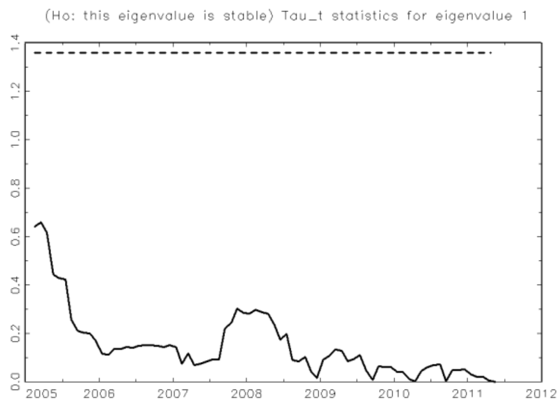
Croatia



Greece



Romania



Turkey

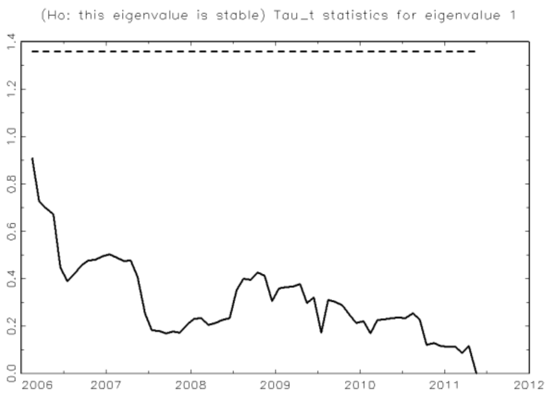


Figure 2. Orthogonal Impulse Responses: Bulgaria

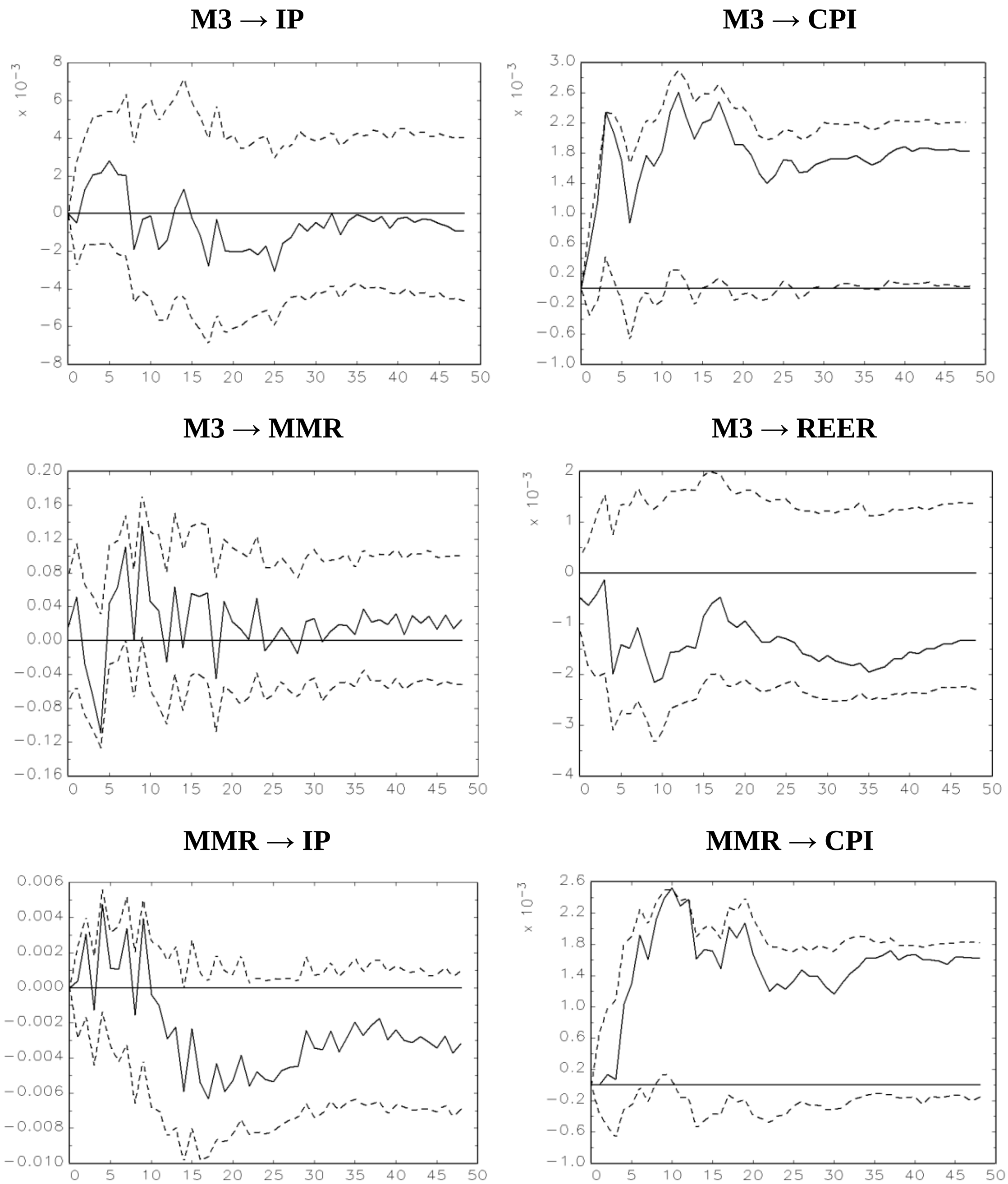
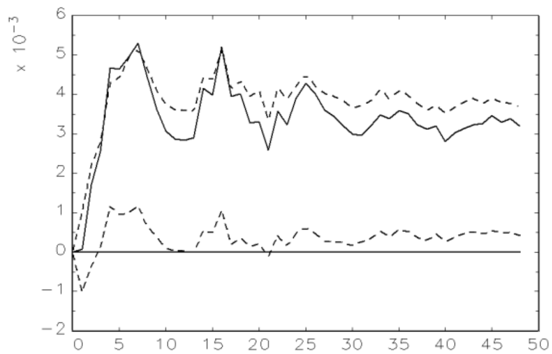
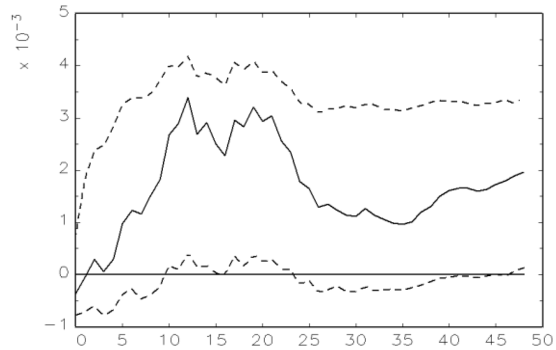


Figure 2 *(continued)*

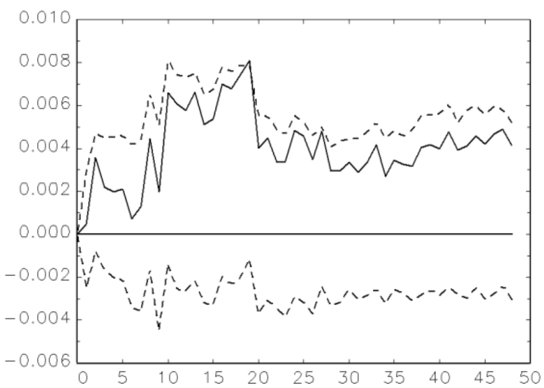
MMR → M3



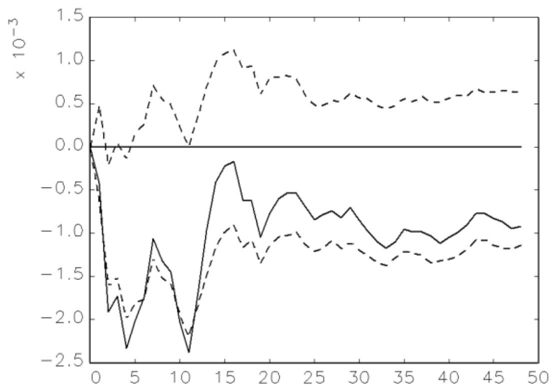
MMR → REER



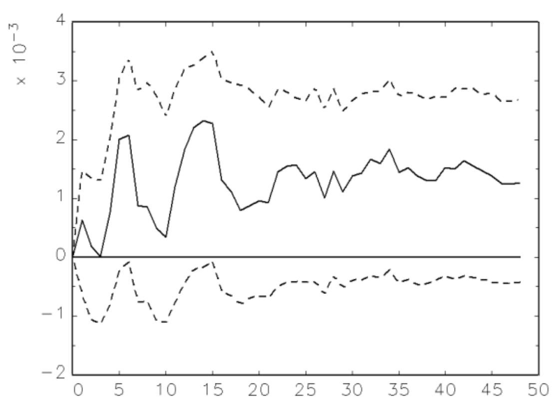
REER → IP



REER → CPI



REER → M3



REER → MMR

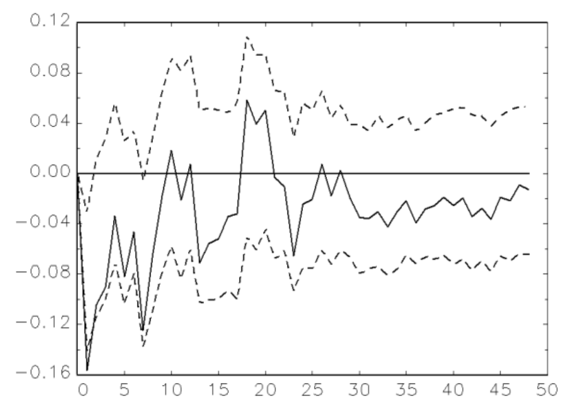


Figure 3. Orthogonal Impulse Responses: Croatia

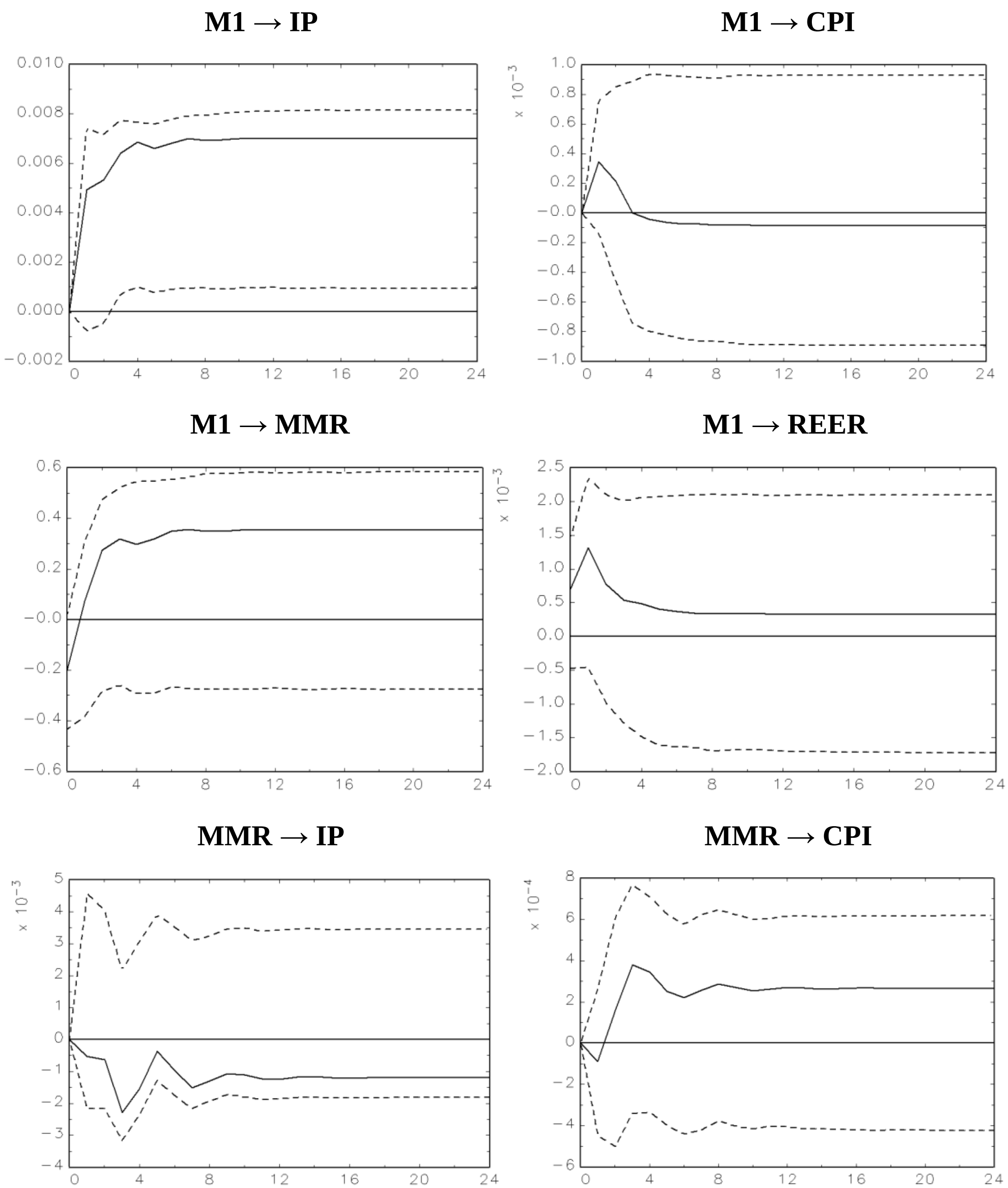


Figure 3 *(continued)*

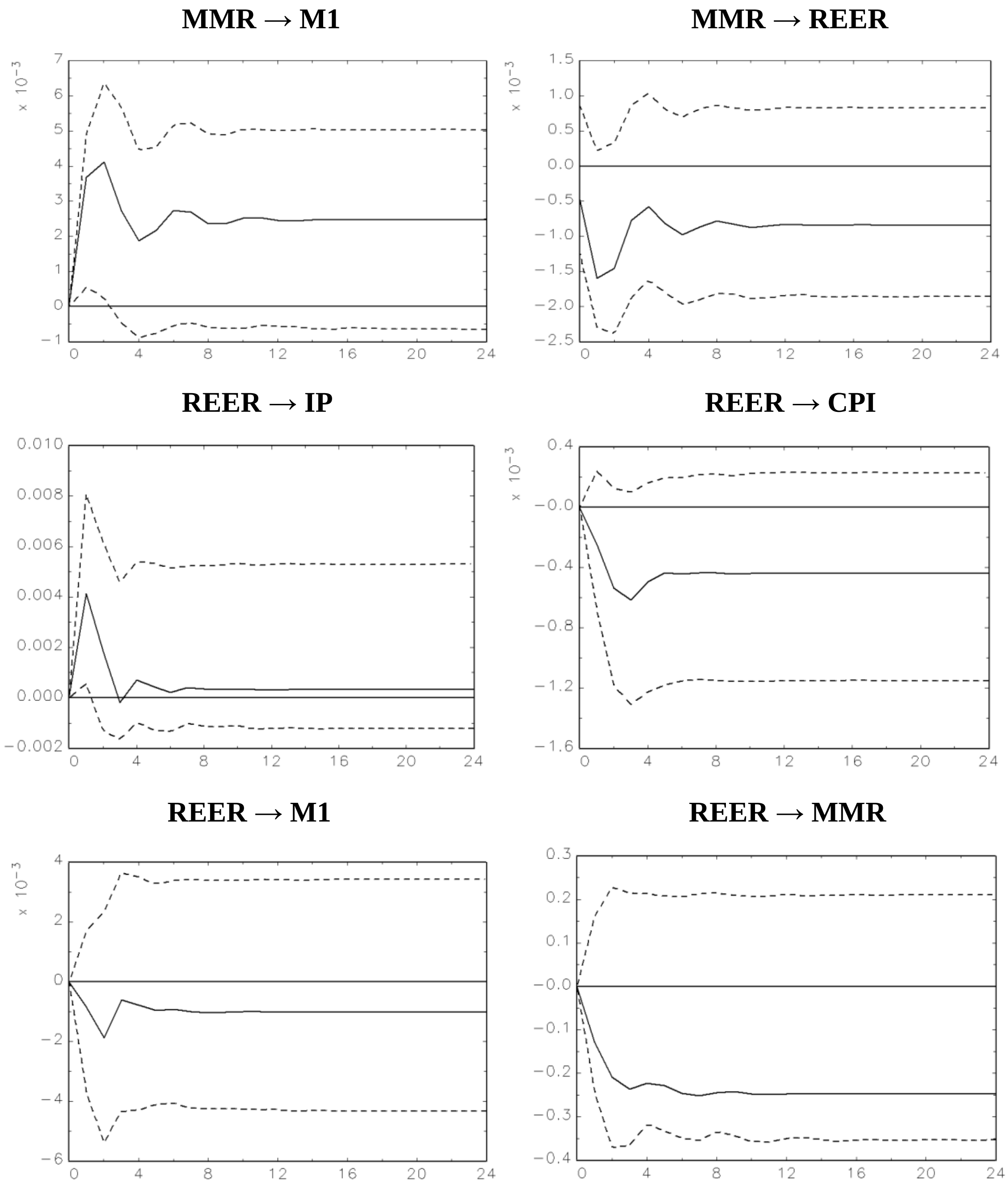


Figure 4. Orthogonal Impulse Responses: Greece

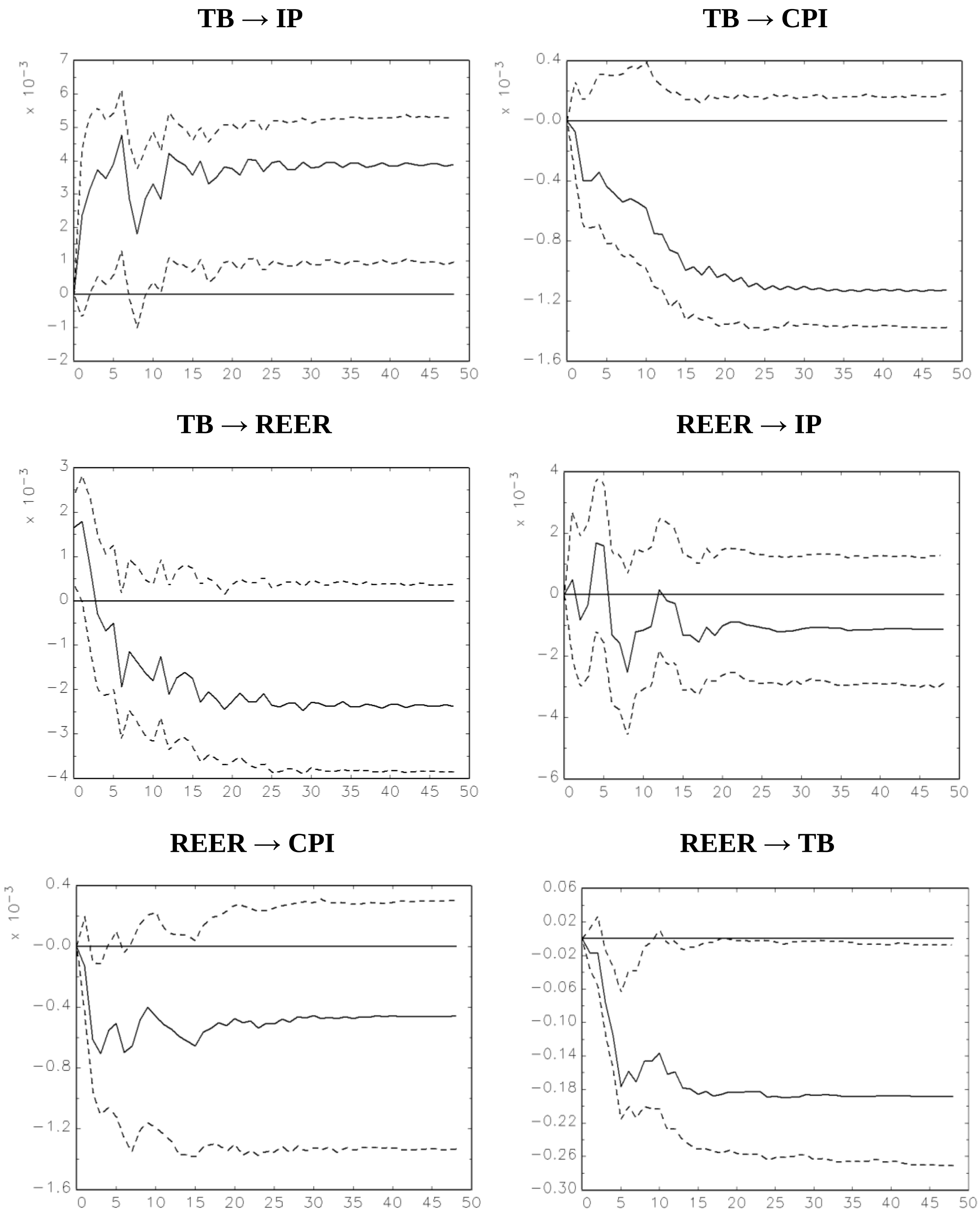


Figure 5. Orthogonal Impulse Responses: Romania

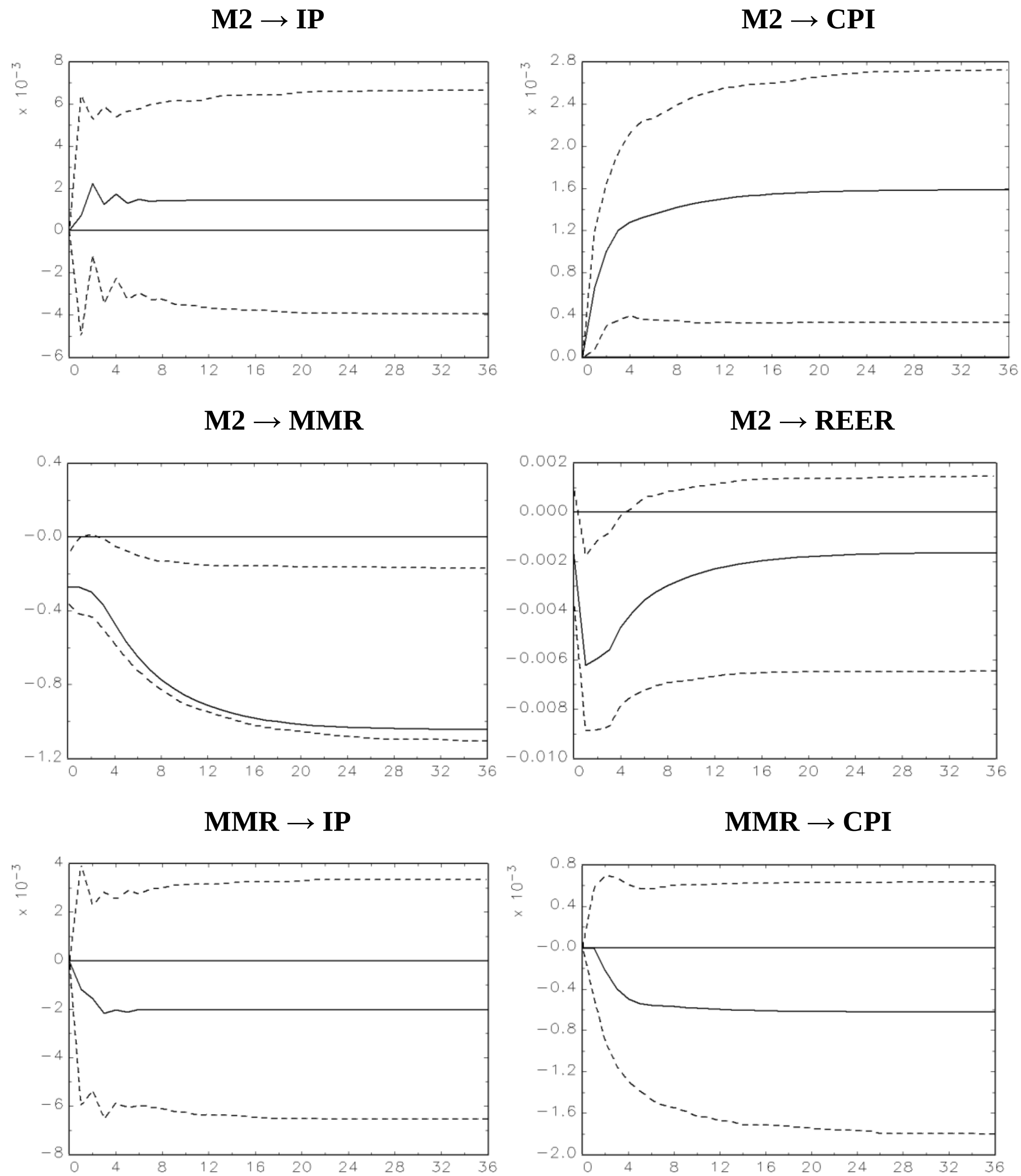


Figure 5 *(continued)*

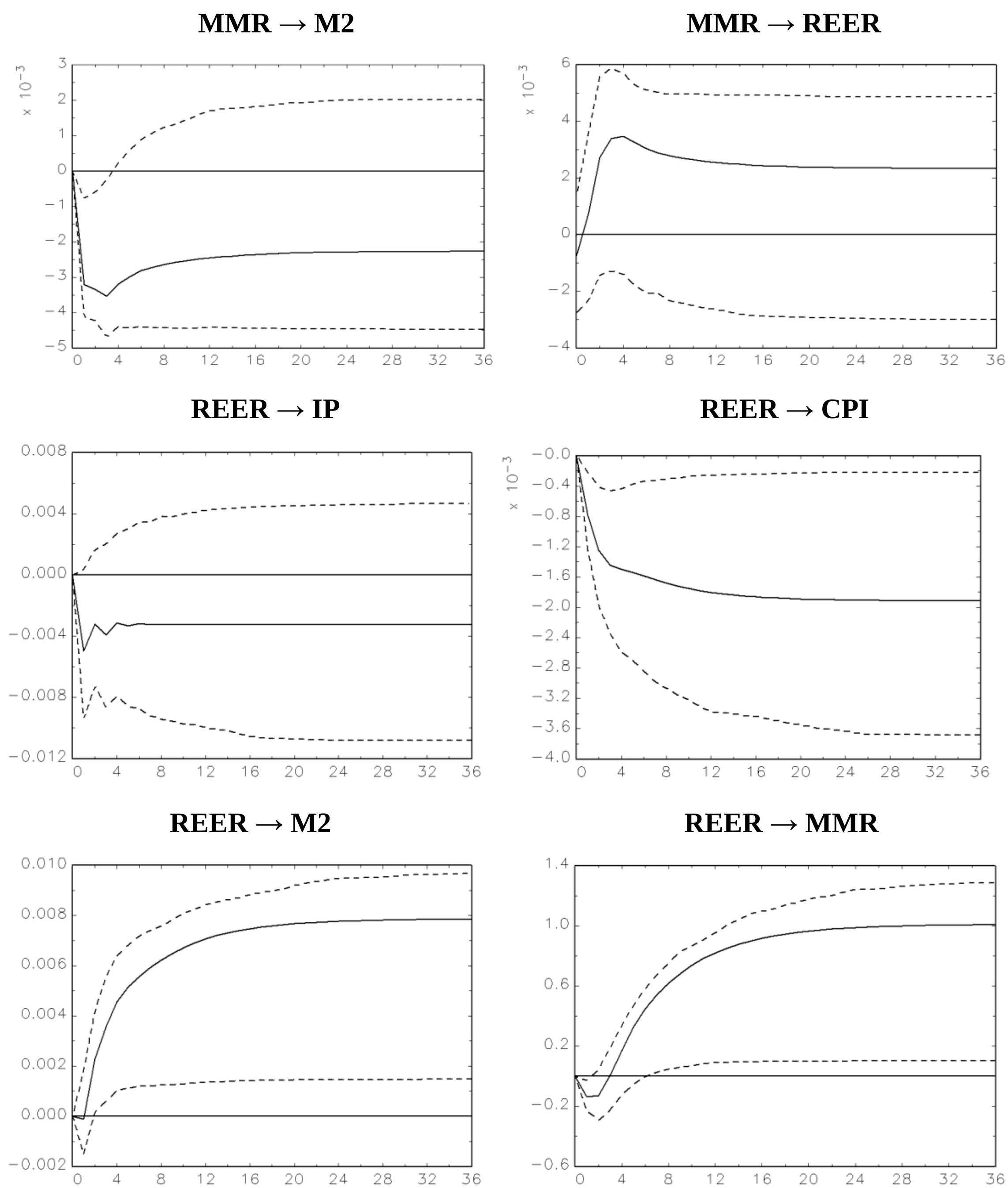
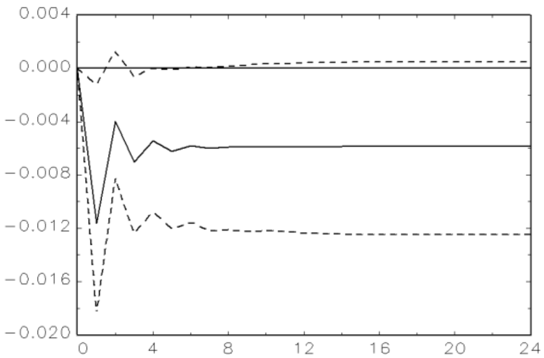
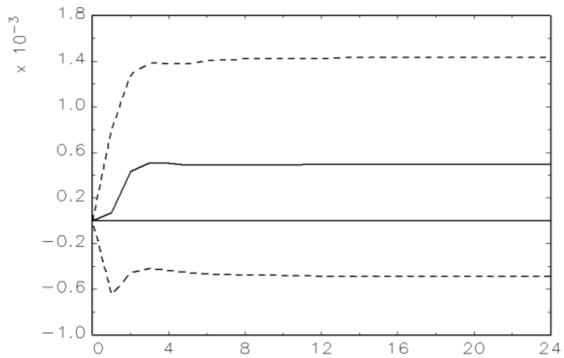


Figure 6. Orthogonal Impulse Responses: Turkey

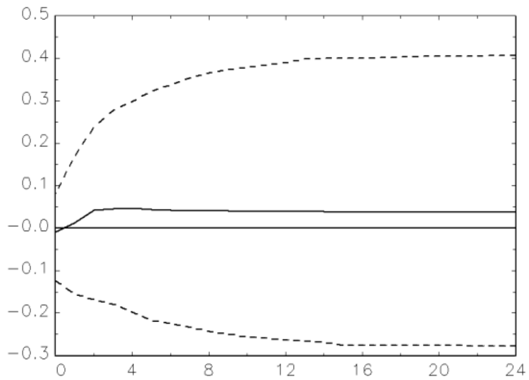
M3 → IP



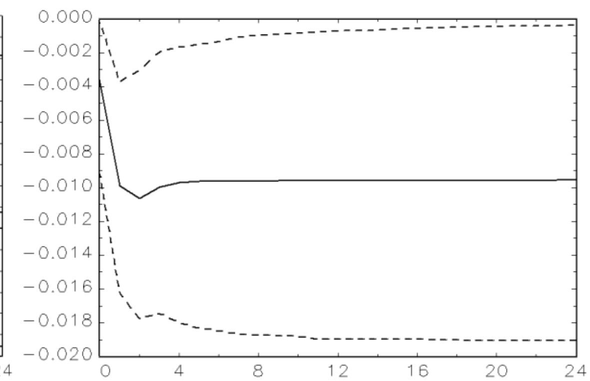
M3 → CPI



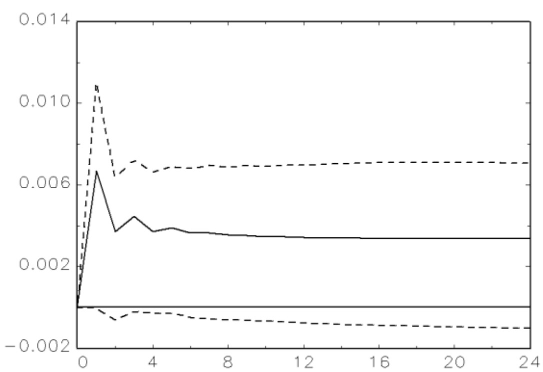
M3 → MMR



M3 → REER



MMR → IP



MMR → CPI

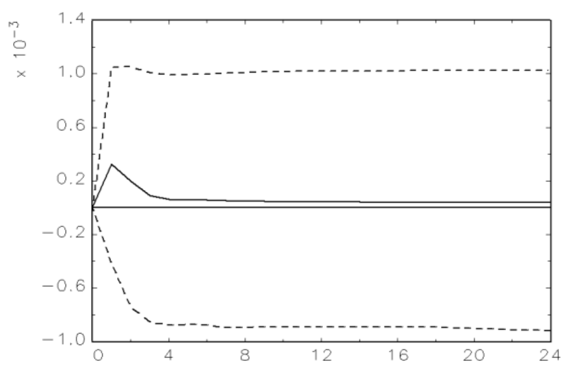
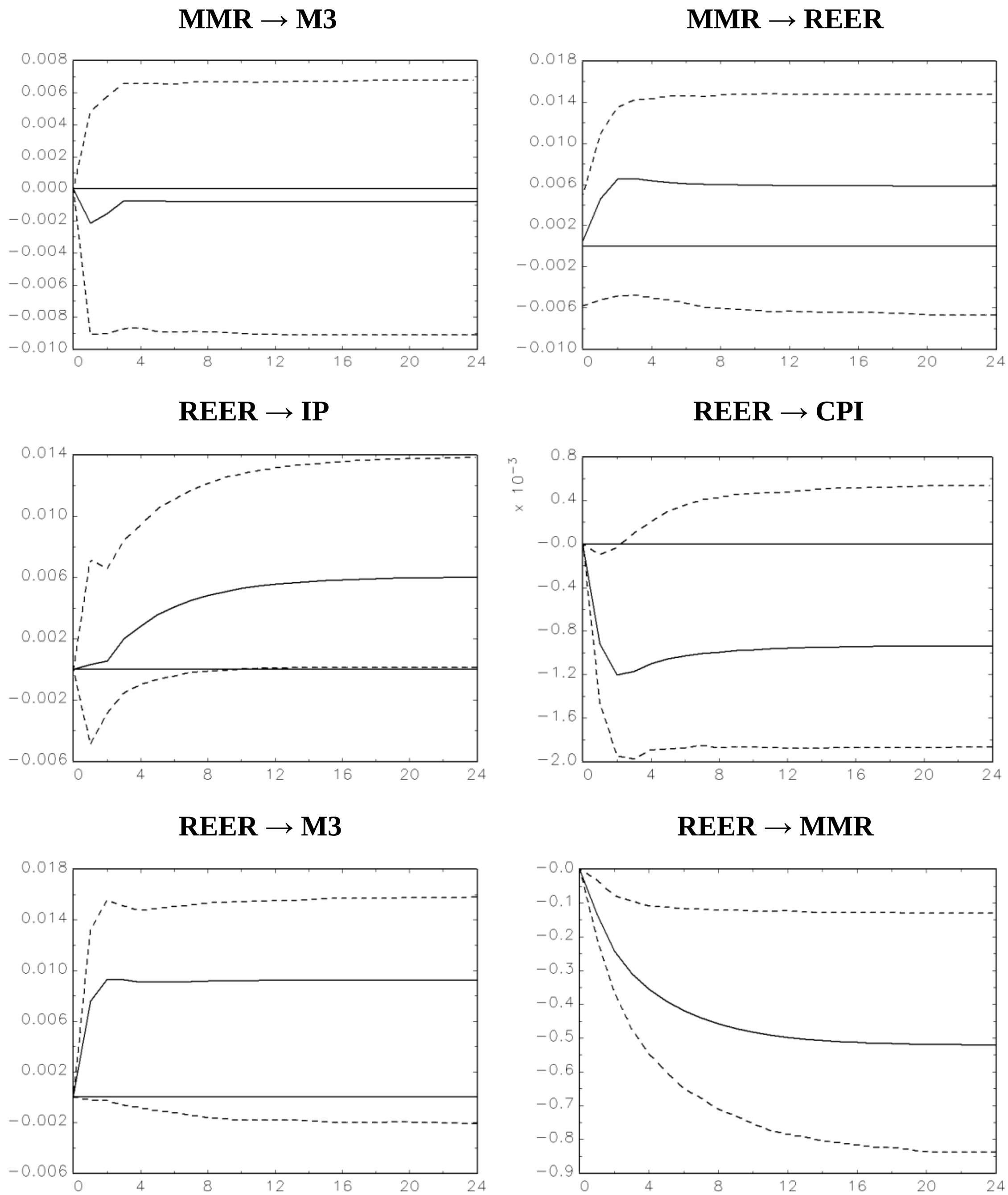


Figure 6 *(continued)*



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