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The Role of the Monetary Policy Stance for the Government Spending Multiplier in Poland

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Abstract

We empirically explore monetary and fiscal policy coordination in Poland. In particular, we study whether the empirical effects of a government spending shock on output depend on the stance of monetary policy. We find no such dependency and conclude after various sensitivity checks, including to slack in the economy, that the government spending multiplier is not dependent on monetary policy or the business cycle. The cumulative multiplier reaches a peak value of 1.11 one year after a government spending shock: a 1 złoty increase in government spending, be it government consumption purchases or government investment or any combination of both, increases real GDP by 1.11 złoty. We identify a crowding-out effect of private investment, but it is relatively small and the overall impact of the government investment shock on GDP is above unity.

JEL Classifications: E62, E63, H50.

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1 Introduction

Recent empirical macroeconomic research on the effects of government spending on economic output has focused on the role of the economic environment in which spending occurs, such as economic slack, interest rates close to zero, and public and private debt relative to GDP. This literature is surveyed in Ramey (2019). Most studies use U.S. data and the literature analysing small open economies such as Poland is sparse. Furthermore, monetary policy may affect fiscal policy not only when interest rates are near zero and monetary policy could potentially undo fiscal policy effects on output. We contribute to this literature by applying state-of-the-art nonlinear econometric methods to study the size of the government spending multiplier in Poland when monetary policy is either tight or loose.[1] The previous nonlinear literature has instead concentrated on the multiplier in relationship to when the economy shows slack or expansion over the business cycle. We also investigate this effect, but the main focus of our analysis is the relationship between the government spending multiplier and the stance of monetary policy. Our quarterly data cover the period from 1999Q1 to 2021Q4. We estimate cumulative multipliers directly with local projections and instrumental variables, following Ramey and Zubairy (2018), among others. In addition, we explore the performance of a COVID-19 stringency index and estimate the government spending multipliers in a sample with and without the COVID-19 period.

The rest of the paper is organized as follows. Section 2 briefly reviews the related literature with special attention to Polish studies. Section 3 discusses the local projection-instrumental variables estimation method, shock identification, weak instrument issues, and alternative definitions of the monetary policy stance. Section 4 presents empirical results for multipliers in regimes of tight and loose monetary policy and also in states of economic slack and expansion. Section 5 explores the robustness of the results and Section 6 concludes.

2 Literature Review

2.1 The Value of the Fiscal Multiplier

Fiscal multipliers measure the short-term output effects of discretionary fiscal policy. They play a key role in planning and forecasting the effects of policy actions. However, the problem is that there is no generally agreed methodology to calculate fiscal multipliers. In addition, it is difficult to isolate the direct effect of fiscal policies on GDP, in particular for tax changes. For example, automatic fiscal

[1] There is a long-standing debate on the interdependence of monetary and fiscal policy. See, for example, Sargent and Wallace (1981), and Davig and Leeper (2011) for a more recent theoretical contribution.

stabilizers lead to increases in government spending in recessions as the government pays benefits to workers being laid off and also tax revenue falls as output and employment fall. Moreover, it is difficult to identify exogenous fiscal shocks, not related to business cycle movements, from observed fiscal outcomes.

Most studies suggest that fiscal multipliers are around unity for government spending increases and about half that size for tax cuts, under normal circumstances. But there is no agreement in the literature on the value of the fiscal multipliers. In some instances, when the fiscal multipliers cannot be reliably estimated due to a lack of available data, one may use the so-called bucket approach (Batini et al. 2014). The method is based on grouping together the countries that are likely to have similar multiplier values based on their characteristics.

Ramey and Zubairy (2018) point out that if the multipliers are indeed below unity, an increase in government spending does not stimulate private activity and, moreover, fiscal consolidations based on a decrease in government purchases are unlikely to harm the private sector. Blanchard and Leigh (2013), on the other hand, argue that multipliers estimated prior to the recent global financial crisis are around 0.5, while multipliers after the crisis are substantially larger. They argue that underestimation of fiscal multipliers early in the crisis contributed significantly to growth forecast errors.

2.2 Nonlinear Effects of Fiscal Policy

Before the world financial crisis of 2008 the fiscal multipliers were often estimated on the whole sample, starting mostly after WWII, and they were linear, in the sense that they were not dependent on the state of the economy. After the crisis much new literature on fiscal multipliers appeared and researchers introduced nonlinearities for the model parameters changing with some chosen threshold or transition variable indicating changes in the state of the economy. A more detailed literature review on state-dependent government spending multipliers may be found, for example, in Haug and Power (2022).

The first group of studies tests how the state of the business cycle affects fiscal multipliers. The most common conclusion here is that the fiscal multipliers are higher during recessions than during expansions (Auerbach and Gorodnichenko 2012; Gechert and Rannenberg 2018). However, the evidence is not conclusive. Empirical studies for the U.S. economy find opposing results for government spending multipliers being larger during recessions than during normal times.[2] The second group of papers tests how monetary policy and private debt affect fiscal multipliers. The government

[2] See, for example, the different results in Barro and Redlick (2011), Auerbach and Gorodnichenko (2012), Ramey and Zubairy (2018) and Fazzari et al. (2015, 2021).

spending multiplier may, or may not, be different during times of very low interest rates near the zero lower bound (e.g., Ramey 2011; Ramey and Zubairy 2018; Bonam et al. 2022), or during times of relatively high private debt (Bernardini and Peersman 2018). Fazzari et al. (2021) consider a model that nests most of the above specifications and find empirical evidence for state-dependent U.S. fiscal multipliers, with government spending multipliers reaching values between 1.15 and 1.23 in periods of excess economic slack.

What is more, fiscal multipliers may vary depending on some structural characteristics of an economy. Multipliers tend to be smaller in more open economies, because the demand-leakage through imports is more pronounced. Also, fiscal multipliers are larger for countries with a more rigid labour market, lower automatic fiscal stabilizers, fixed exchange rates, lower debt levels, less difficulties to collect taxes and less expenditure inefficiencies (Batini et al. 2014). Thus, there are also some empirical papers that test fiscal multiplier nonlinearities depending on, for example, the level of private debt (Bernardini and Peersman 2018), the stress level in financial markets (Afonso et al. 2018), the composition of the fiscal adjustment (Gechert and Will 2012), and the openness of the economy (OECD 2009). Finally, there is evidence that fiscal multipliers depend on the fiscal position, particularly the output cost of consolidation can be lower in times of weak public finances (Huidrom et al. 2020).

2.3 Specification Choices in Calculating Multipliers

There are two main methods to derive fiscal multipliers: empirical estimation using structural vector-autoregression (SVAR) models and dynamic stochastic general equilibrium (DSGE) model-based approaches. The pros and cons of each method may be found in Batini et al. (2014). In this paper we concentrate on SVAR-based models that rely on country specific data.

Čapek, and Cuaresma (2020) provide a study on the role played by data and specification choices for the estimated fiscal multipliers in SVAR models. The authors use quarterly data for European countries from 1999 to 2014. They show that these choices (i.e., the choice of the deflator, the composition of government spending and government revenue, the choice of the structural shock identification strategy, and the number of variables) can have an important impact on the precision of multiplier estimates leading to a cumulative change in the multiplier of as much as 0.4. Interestingly, using the harmonised index of consumer prices instead of the GDP deflator to deflate nominal variables, or using the EU's ESA 95 instead of ESA 2010 data definition can lead to a 0.12 increase in spending multipliers. Poland is included in this study in a group of Eastern EU countries and all the results apply to Poland as well.

2.4 Fiscal Multipliers for Poland

There is a small but growing literature on fiscal multipliers for Poland (e.g., Łaski et al. 2012; Bencik 2014; Baranowski et al. 2016; Haug et al. 2019; Szymańska 2019). Łaski et al. (2012) compare the factors of growth in Poland and the Czech Republic in 2009. They calculate fiscal multipliers, among others, using a purely accounting-based method. The authors use the demand side approach, where GDP is obtained as a weighted sum of government spending, private investment and exports. Batini et al. (2014) note that the method may lead to inaccurate estimates of the overall impact of fiscal measures, as second round effects are difficult to quantify.

Łaski et al. (2012) report the fiscal multiplier for Poland equal to 1.6 in 2009 (1.56 in 2008), and underline the striking difference with the multiplier for the Czech Republic equal to 1.25 in 2009 (1.14 in 2008). Haug et al. (2019) estimate a linear SVAR model for Poland using quarterly data between 1998Q1 and 2013Q3. For the identification of structural shocks they impose contemporaneous restrictions based on calculated elasticities, following Blanchard and Perotti (2002). They report an initial government spending multiplier of 0.70, which later peaks at 1.61 for the cumulative multiplier.

Another group of papers concentrates on fiscal multipliers in so-called Visegrád Four (V4) countries – Poland, the Czech Republic, Slovakia, and Hungary (e.g., Bencik 2014; Baranowski et al. 2016; Szymańska 2019). These countries are not only linked by being neighbours, but also they are still in the process of converging towards advanced European economies. Baranowski et al. (2016) analyse the effectiveness of fiscal policy in the V4 countries. They estimate a threshold-switching VAR model with output, net taxes, and government spending, using quarterly data between 2000Q1 and 2012Q4. They apply Blanchard and Perotti's (2002) identification scheme, but assume that the tax shock is preceding the government spending shock. Unfortunately, they do not report the value of fiscal multipliers for separate economies. Also, they report only instantaneous elasticities of output as the fiscal multipliers (their Table 2). The reader can assume (based on their Figure 1) that the government spending multiplier is positive in recessions and negative in expansions.

Szymańska (2019) concentrates on government spending shocks in the same group of countries. Szymańska uses the three-variable SVAR model as in Baranowski et al. (2016) and calculates the fiscal multipliers at various horizons after a shock but does not consider any nonlinearities. Szymańska analyses the time period from 2002Q1 to 2018Q1. Szymańska reports the cumulative government spending multiplier for Poland equal to 1.46 after one year and the peak multiplier is equal to 1.76 after 7 quarters. Szymańska reports some sensitivity of the multipliers to using different tax measures and different lag structures.

On the other hand, Bencik (2014) applies a simplified smooth-transition SVAR model to calculate government spending multipliers during recessions and expansions in Poland, the Czech Republic, Hungary, and Slovakia. The author shows higher fiscal multipliers in recessions than in expansions for V4 countries, but Bencik does not report the results for separate economies.

Additionally, there is a group of papers on the output costs of fiscal consolidation (e.g., IMF 2012, Blanchard and Leigh 2013) and some of them consider the Polish economy as one of the panel countries (Cugnasca and Rother 2015; Górnicka et al. 2018). Cugnasca and Rother (2015) estimate the short-term output cost of fiscal consolidation for a panel of EU countries between 2004 and 2013. They concentrate on the Excessive Deficit Procedure (EDP) recommendation and use it as an exogenous instrument for discretionary fiscal policy actions. Their results show that a 1% of GDP improvement in the structural primary government budget balance leads to a 0.50 % decrease in GDP in EU countries and a 0.76 % decrease in countries with the currency pegged to the euro. Moreover, they test how the results depend on the state of the business cycle, trade openness, composition of fiscal adjustment, and the degree of stress in credit markets but do not report results for individual countries.

Górnicka et al. (2018) provide a similar study. They analyse EU countries which were subject to EDPs between 2009 and 2015. Their findings show that fiscal multipliers (defined as the ratio between output gaps and structural budget balances) increased over time – from about 0.25 in the early years of the crisis to about 0.66 in the later years. However, they do not confirm the hypothesis that ex-post fiscal multipliers have been substantially above 1 during the crisis, in contrast to Blanchard and Leigh (2013). Górnicka et al. (2018) report the value of the multiplier for Poland as equal to -0.1 (for 2009 - 2011) and 0.7 (for 2012 – 2015).

Summing up, the literature offers various approaches to measure fiscal multipliers and the results from different approaches are not directly comparable with each other. As far as the fiscal multipliers in Poland are concerned, empirical studies seem to suggest that they are on a moderate level. It corresponds to the evidence based on the bucket approach, which suggests that the first year overall multiplier for Poland is in the range of 0.4 to 0.6 (Batini et al. 2014).

3 Methodology

We apply the local projection (LP) method of Jordà's (2005) to estimate the government spending multipliers and the associated impulse response functions. Following the recent literature, we focus on cumulative multipliers and impulse responses, in particular we follow Bernardini and Peersman (2018) and Ramey and Zubairy (2018). These authors explore the role of private debt and the state of the economy over the business cycle for the size of the government spending multiplier for

U.S. data, respectively. The nonlinear models applied in their study have two states (high and low private debt states, and expansions and recessions) in an SVAR-based framework. Instrumental variable (IV) methods are used in connection with LP to estimate the cumulative multipliers and impulse responses directly.

3.1 Government Spending Multipliers and Local Projections

Ramey (2019) points to a general problem when a model is specified with variables in logarithms and the parameter estimates represent elasticities. In order to get proper multipliers in dollar-for-dollar terms, the elasticities need to be evaluated with a conversion factor. The conversion factor is generally specified as the ratio of the sample average of GDP to the sample average of government spending. However, this ratio often varies considerably over the sample period and changes with the length of the time series considered. In order to overcome these problems, Ramey and Zubairy (2018) follow a transformation proposed by Gordon and Krenn (2010).[3] Variables are specified in levels, except for real GDP and real government spending, which are divided by the stochastic trend of real GDP. This transformation produces multipliers that give the response of real GDP in currency units to a one currency unit shock to government spending. In the first instance we apply the Hodrick-Prescott (HP) filter to extract the stochastic trend from the real GDP series (Hodrick and Prescott, 1997). We set the smoothing parameter λ equal to 1,600, as is standard for quarterly data.

Following Ramey and Zubairy (2018), we specify for each horizon h the following nonlinear model:

$$y_{t+h} = I_{t-1}\left[\alpha_{A,h} + \psi_{A,h}(L)z_{t-1} + \beta_{A,h}shock_t\right] + (1 - I_{t-1})\left[\alpha_{B,h} + \psi_{B,h}(L)z_{t-1} + \beta_{B,h}shock_t\right] + \varepsilon_{t+h}, \quad h = 0, 1, \dots, 12, \tag{1}$$

where y_{t+h} is real GDP in period $t+h$ divided by its own stochastic trend. The indicator variable, I_{t-1}, picks the state of the economy at time t-1 with a value of 0 for periods of loose monetary policy (or slack when the business cycle is studied), labelled state *A*, and a value of 1 for periods of tight monetary policy (or expansions), labelled state *B.* As in Bernardini and Peersman (2018) and Ramey and Zubairy (2018), the state is determined by the state in the period before the shock hits to avoid contemporaneous feedback from government policy to the state. The variable $shock_t$ is a government spending shock at time *t*, divided by the stochastic trend of real GDP. The shock is uncorrelated with the error term ε_{t+h}. Hence, $\beta_{A,h}$ and $\beta_{B,h}$ are the non-cumulative responses of real GDP to the

[3] Bernardini and Peersman (2018) use an alternative transformation that does not rely on extracting a stochastic trend. The changes in real GDP and in government spending are scaled instead by y_{t-1}.

government spending shock for tight and loose monetary policy, or economic slack and expansion, respectively. The parameter α is a constant, $\psi(L)$ is a lag polynomial, and z_{t-1} is a vector of lagged control variables, which includes lags on y_t. The LP method of Jordà's (2005) differs from the conventional method of estimating impulse response functions. The conventional method instead estimates parameters only at horizon zero and iterates forward to derive the effects on y_{t+h}, without estimating a new set of parameters for each horizon.

3.1.1 The Cumulative Multiplier

As in Ramey and Zubairy (2018), we directly estimate the cumulative multiplier. We estimate the following equation with the local projection – instrumental variables (LP-IV) method:

$$\sum_{j=0}^{h} y_{t+j} = I_{t-1}\left[\gamma_{A,h} + \phi_{A,h}(L) z_{t-1} + m_{A,h}\sum_{j=0}^{h} g_{t+j}\right] + (1 - I_{t-1})\left[\gamma_{B,h} + \phi_{B,h}(L) z_{t-1} + m_{B,h}\sum_{j=0}^{h} g_{t+j}\right] + \omega_{t+h},$$
$$h = 0, 1, \dots, 12, \qquad (2)$$

where $I_{t-1}shock_t$ and $(1 - I_{t-1})shock_t$ are used as instruments for cumulative government spending $\sum_{j=0}^{h} g_{t+j}$ in the corresponding state. The estimated cumulative spending multipliers at horizon h for state A and B are the estimates of $m_{A,h}$ and $m_{B,h}$.

The advantage of the IV method is that the cumulative multipliers and their standard errors are estimated directly from equation (2). Also, the shock and the government spending variable are both allowed to exhibit measurement error, as long as their measurement errors are uncorrelated. The LP-IV regressions are implemented with the two-stage least-squares Stata command *ivreg* (Baum et al. 2010). To test for autocorrelation in the presence of IVs, we apply the *ivactest* test suggested by Baum et al. (2007). This test indicates serial correlation. Hence, we run all IV regressions with the Newey-West correction. The options *bw(auto)* and *robust* were selected to calculate heteroskedasticity and autocorrelation consistent (HAC) variance estimates, based on a Bartlett kernel and an associated automatic bandwidth selection criterion (Baum et al. 2007).

3.1.2 Shock Identification

We identify government spending shocks, referred to as Blanchard-Perotti shocks, following, inter alia, Bernardini and Peersman (2018) and Ramey and Zubairy (2018).[4] It is assumed that government spending does not respond contemporaneously to other structural shocks. This is a reasonable assumption for the identification of structural government spending shocks, because there are no automatic stabilisers affecting government purchases, as is the case for taxes, and government spending changes are lagged due to the time it takes for making and implementing government spending decisions.[5] We assume that the government spending policy is backward-looking and evolves as:

$$g_t = \psi(L)z_{t-1} + shock_t, \tag{3}$$

where g_t is determined by the same set of lagged control variables used in equation (1) above, z_{t-1}, and an orthogonal shock ($shock_t$) of autonomous changes to government spending.

3.1.3 Weak Instruments and Hypothesis Testing

With LP-IV estimation, a concern about the quality of the instruments arises. Using weak IVs may lead to biased and unreliable estimates (Andrews et al. 2019). To detect weak instruments, we calculate for the first stage estimates, in the two-stage-least-squares procedure of the LP-IV method, the Kleibergen-Paap rk F-statistic (Kleibergen and Paap 2006). This F-statistic is calculated for each state and horizon when estimating equation (2). A rule of thumb for linear IV regressions is an F-statistics below 10 indicating a weak instrument (Staiger and Stock 1997, and Stock and Yogo 2005). In the presence of first stage regression errors that are heteroskedastic and serially correlated, Montiel Olea and Pflueger (2013) propose a stricter threshold.[6] The value of their threshold at a 10% significance level has a value of 23.1.[7] Following Mertens and Montiel Olea (2018), we choose an F-statistic of 10 or above to indicate that an instrument is not weak. For values above 23.1, we refer to an instrument as being particularly strong.

We test whether the cumulative multiplier is statistically significantly different across tight and loose monetary policy states and, alternatively, across business cycle states. The null hypothesis is that the cumulative multipliers are the same in both monetary policy or business cycle states for a given

[4] These authors also explore large U.S. military spending shocks due to events unrelated to the U.S. business cycle. These shocks are identified narratively. This approach seems not suitable for Poland.
[5] In an SVAR with a recursive Cholesky ordering, this would amount to putting government spending first.
[6] They consider a linear model with one endogenous regressor only. See also Montiel Olea et al. (2021).
[7] The 10% significance level is chosen because of the relatively small sample size. Montiel Olea and Pflueger's (2013) threshold values were generated with their Stata code *weakivtest* (Pflueger and Wang 2013, 2015) where $\tau = 10\%$. The 5% level test has a critical value of 37.4.

horizon. In this case, the model is linear and multipliers are not affected by the stance of monetary policy or the business cycle state:

$$\sum_{j=0}^{h} y_{t+j} = \gamma_h + \phi_h(L)z_{t-1} + m_h \sum_{j=0}^{h} g_{t+j} + \omega_{t+h}, \quad h = 0, 1, \ldots, 12 \tag{4}$$

The alternative hypothesis is that the multipliers are different across states. In this case, government spending multipliers are dependent on whether monetary policy is tight or loose, and for the business cycle on whether the economy is in a recession or expansion. In our tables, we report for this hypothesis test the HAC-based p-values when instruments are not weak, and Anderson and Rubin's (1949) test (AR) p-values when instruments are weak, because the AR test is robust to weak instruments (but possibly less powerful than the HAC-based test). The AR test p-values are reported only when either or both states have a Kleibergen-Paap rk F-statistic value below 10.

3.1.4 Lag Choice, Unit Roots and Cointegration

The lag length for the estimation of equation (2) is chosen based on Akaike's information criterion and on Schwarz's Bayesian information criterion. The two criteria mostly choose the same number of four lags for the control variables in the vector z_t in all our regressions below, unless indicated otherwise. In regards to nonstationary variables, we test for unit roots and cointegration and find evidence in favour of both. Hence, a specification with variables in first differences is not appropriate, whereas an error-correction model could be considered. However, a Monte Carlo study by Gospodinov et al. (2013) finds that if the exact magnitude of roots is unknown or unclear, conventional linear VAR-based impulse responses with a VAR in levels are more robust than those from a VAR that imposes cointegration and unit roots.[8] Furthermore, Montiel Olea and Plagborg-Møller (2021) show that inference in lag-augmented linear LP-based IRFs is valid with stationary and nonstationary data. We use lag-augmented LP regressions below, i.e., the control variables in z_t include lags on the dependent variable y_t, and we do not impose cointegration.

3.2 Defining States of Tight and Loose Monetary Policy

To define the states of tight and loose monetary policy we apply three different approaches, based either on the deviation of the real interest rate from the trend, developments in Monetary Condition Indices or the deviation of the nominal interest rate from its value implied by the monetary policy rules.

[8] Linear LP-based and linear conventional IRFs are the same in VAR models with unrestricted lag structures, i.e., as the sample size goes to infinity (Plagborg-Møller and Wolf 2021).

3.2.1 An Indicator Variables Based on the Real Interest Rate

In order to identify the periods of tight and loose monetary policy for the indicator variable, I_{t-1}, in equations (1) and (2) we appeal to a theoretical model developed by Davig and Leeper (2011). They consider an active monetary policy as a policy that reacts to fiscal expansion and an associated increase in inflation by raising the nominal policy interest rate. This leads to an increase in the real interest rate and thus reduces consumption and investment demand. It dampens the stimulative effects of the expansionary fiscal policy and with it the fiscal multiplier. On the other hand, a passive monetary policy does not respond with interest rate increases to inflation and the impact of expansionary fiscal policy is amplified. By not pushing up the nominal interest rate, monetary policy allows higher current and expected inflation to translate into lower real interest rates than otherwise.

In order to make this approach operational, we assign to the indicator variable a value of 0 when the 3-month ex-post real WIBOR interest rate is equal to or below its stochastic trend, which we define as a loose monetary policy state.[9] When the real WIBOR rate is above its stochastic trend in a given quarter, we define the state of that quarter as tight monetary policy and assign a value of 1 to the indicator variable. To extract the stochastic trend form the real WIBOR variable, we apply again the HP filter with a smoothing parameter of 1,600.[10] Over our sample period, the indicator variable seems to perform quite well in term of picking relative peaks and troughs of the real interest rate. Among the 92 quarters in our sample, 46 are identified as having loose monetary policy and 46 as having tight monetary policy (Figure 1).

3.2.2 An Indicator Variables Based on a Monetary Condition Index

The restrictiveness of monetary policy is not only described by the real interest rate. Therefore an alternative proxy for loose or tight monetary policy is based on so-called Monetary Condition Indices (MCIs; e.g., Ericsson et al. 1998), developed in the early 1990's. Measuring the monetary stance, MCIs attempt to capture two main transmission channels of monetary policy, the interest rate channel and the exchange rate channel. MCI is calculated as a weighted average of the real interest rate and the real exchange rate relative to its value in a base period. The increase of the MCI indicates an increase in the monetary policy restrictiveness, while its decrease corresponds to monetary policy loosening. MCIs, introduced by the Bank of Canada, were used by central banks as indicators of

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[9] We use the ex-post real interest rate with inflation based on the GDP-deflator. We tried instead ex-ante real interest rates based on a measure of expected inflation and found partial (at three horizons) support for a nonlinear model but the nonlinear impulse responses made little economic sense.

[10] In the robustness analysis in Section 5.2 we apply alternatively the Hamilton (2018) filter.

monetary conditions and also as operational short-run targets for monetary policy. They were also used in the analysis of monetary policy transmission (e.g., Altavilla and Landolfo 2005).

In the present study we calculate eight different measures of MCIs. For both the real interest rate and the real effective exchange rate the base period is 1999Q1. As far as the real interest rate is concerned, we use either the ex-ante real interest rate or the ex-post real interest rate. In this way we capture backward- and forward-looking behaviour both by the central banks and the other economic agents. In addition, in both cases we use either CPI inflation or core inflation (excluding food and energy) as the deflator. For the real effective exchange rate, based on bilateral exchange rates of Polish złoty (PLN) against the euro (EUR) and US dollar (USD), we apply different weights for both bilateral exchange rates: 0.9 for EUR/PLN and 0.1 for USD/PLN, or 0.8 for EUR/PLN and 0.2 for USD/PLN. In calculating all versions of the Monetary Condition Index the interest rate weight is twice as large as the weight on the exchange rate. It is broadly consistent with the relative importance of interest rate and exchange rate channels in Poland in regards to the horizons of the maximum impact of monetary policy for CPI inflation (Chmielewski et al. 2020).

Figure 2 presents MCIs and their median. The MCIs calculated are highly correlated with each other. Based on their median value we identify 56 quarters with loose monetary policy and 36 quarters with tight monetary policy in our sample. Longer periods of tightening of monetary conditions are 1999Q2-2001Q2, 2004Q2-2005Q1 and 2007Q2-2008Q3, while longer periods of loosening of monetary conditions are 2002Q1-2004Q1 and 2015Q3-2016Q2. These periods correspond to the periods identified on the basis of the rule-of-thumb applied, according to which a given period is characterized by tightening or loosening of monetary conditions if at least ¾ of the MCIs indicate this direction.

Nowadays MCI measures are less popular than in the early stage of the introduction of formal inflation targeting at central banks.[11] It is due to the fact that they display some weaknesses, in particular they assume constant weights of the interest rate and the exchange rate, along with exogeneity of policy instruments and other variables (e.g., Ericsson et al. 1998; Batini and Turnbull 2002).

3.2.3 An Indicator Variables Based on a Monetary Policy Rule

The third approach to defining the indicator variable is based on the estimation of monetary policy rules (e.g., Taylor 1993, Taylor 2007). It compares the actual short-term interest rate in Poland with the theoretical one, consistent with the estimated behaviour of monetary policy makers. The

[11] For an overview and discussion of problems of the MCI see Osborne-Kinch and Holton (2010).

periods of excessively tight and loose monetary policy are given by positive or negative deviations of the actual interest rate from the interest rate resulting from the monetary policy rule.

To comprehensively assess periods of tight and loose monetary policy we estimate different versions of the response functions of the monetary authority. First, we apply both the conventional Taylor rule (Taylor 1993) as well as its modified version, capturing interest rate smoothing by central banks (Taylor 2007). Second, we consider different proxies for the main explanatory variables in the Taylor rule, i.e., for inflation and economic slack (the output gap). In the former case, we apply either CPI inflation or core inflation. We derive the measure of economic slack by applying alternatively the HP, Christiano-Fitzgerald (2003) (CF) and Hamilton (2018) filters. Third, we take into consideration that the central bank may be responsive either to current, past or future (projected) deviations of inflation from the inflation target, i.e., it can be either backward- or forward-looking. Therefore the general form of a monetary policy rule based on the conventional Taylor rule is the following:

$$i_t = \alpha_0 + \alpha_1\left(\pi_{i,t+5-j} - \pi^*\right) + \alpha_2 \hat{y}_{k,t} + \varepsilon_t, \tag{5}$$

while the specification of the Taylor rule with interest rate smoothing is

$$i_t = \beta_0 + \beta_1 i_{t-1} + (1-\beta_1)\left[\beta_3\left(\pi_{i,t+5-j} - \pi^*\right) + \beta_4 \hat{y}_{k,t}\right] + u_t, \tag{6}$$

where i_t is the short-term nominal interest rate (3-month WIBOR) at time t, $\pi_{i,t}$ stands for the i-th measure of inflation (CPI inflation or core inflation), while $\hat{y}_{k,t}$ is the k-th measure of the economic slack (obtained using either the HP, CF or Hamilton filter). The parameter $j = \{1, 2, \ldots, 6\}$ is used to define the lead or lag with which the deviation of inflation from the Narodowy Bank Polski (NBP, central bank of Poland) inflation target enters the monetary policy rule. The extreme cases we analyse are the forward-looking central bank that considers the inflation gap four quarters ahead, which is broadly consistent with the lags in the monetary transmission mechanism in Poland (e.g., Chmielewski et al. 2020) and the backward-looking central bank that pays attention to the deviation of inflation from the target in the previous quarter.

Estimation results suggest that in terms of empirical fit the conventional Taylor rules are substantially inferior to the monetary policy rules capturing the interest rate smoothing behaviour. It is fully in line with our intuition given that interest rate smoothing by central banks is a well-documented stylised fact. The median coefficient of determination (adj. R^2) from the former models is 0.59, while in the case of the latter ones it is close to 0.98. The ranges of these coefficients from both types of models do not overlap each other – they are 0.46-0.86 and 0.96-0.99, respectively. Therefore in defining periods of tight and loose monetary policy we rely on the results from the interest-rate-smoothing augmented Taylor rules. Figure 3 presents the interest rate gaps based on the above models, defined as differences between the actual short-term interest rate (3-month WIBOR) and the theoretical one, consistent with the usual conduct of monetary policy. We can note that these deviations are rather small. Longer periods of excessively tight monetary policy based on the median interest rate gap are 2001Q1-2003Q3 and 2018Q2-2021Q3, while longer periods of excessively loose monetary policy are 1993Q3-2000Q4 and 2013Q3-2017Q3. For the median case we identify 46 quarters with monetary policy characterised as loose and 46 quarters as tight. The indicator variable based on the real interest rate picked the same number of quarters with loose and tight monetary policy, however, the periods do not all overlap and several periods start and/or end in different quarters.

In conceptual terms monetary policy rules and Monetary Condition Indices highlight different aspects of analysing the monetary policy stance. Analysis based on monetary policy reaction functions shows whether the actual short-term interest rate is above or below its value consistent with the long-run behaviour of the monetary authority. MCIs are aimed at assessing the direction of change in monetary conditions, without comparing their values to a specific benchmark. Therefore it is natural that for only 60% the two measures show the same regime.[12] It is due to the fact that the periods of the short-term interest rate being above or below the benchmark value can comprise episodes of monetary policy tightening and loosening, in the same way as a positive output gap can be associated with deteriorating or improving real GDP.

4 Results

Quarterly data used are described in Appendix A. The sample period covered is from the first quarter of 1999 to the fourth quarter of 2021. The start date is due to data availability and the end date reflects the latest data available when this research was started. The baseline model includes the following seven variables:

a. in real terms and in levels:

- GDP
- government purchases (government spending on consumption and investment)
- net tax revenue (net of government transfers and subsidies)
- effective exchange rate;

b. in nominal terms and in percentages:

- GDP-deflator based inflation rate
- 3-month WIBOR interest rate;

c. a scaled dummy variable to account for the COVID-19 period:

[12] Detailed results for all monetary policy rule and MCI estimations are available from the authors on request.

- the COVID-policy stringency composite index takes values from 0 to 1, with 1 indicting the most restrictive policy in terms of workplace closures, travel bans, etc.

The choice of variables for inclusion in the baseline model is guided by related recent empirical studies combining monetary and fiscal policy in a small open economy model, such as Dungey and Fry (2009) for New Zealand, and Haug et al. (2019) for Poland.[13] Due to our relatively small sample size, we choose a parsimonious specification in terms of variables included that still delivers a robust model for combining monetary and fiscal policy analysis.[14] The first six variables above form the vector of control variables, z_t, used in equations (1) to (4), along with the COVID-19 dummy variable.

We estimate equation (2) by LP-IV and define the indicator variable, I_{t-1}, which takes values of 0 and 1, to capture loose and tight monetary policy and, alternatively, economic slack and economic expansion, as defined below in Section 4.4. We present results for the baseline model with potential nonlinearities introduced sequentially by three alternative definitions of the monetary policy stance and by economic slack. All multipliers reported are cumulative złoty-for-złoty multipliers.

4.1 Results with the Monetary Policy Stance Based on the Real WIBOR Interest Rate

Table 1 reports results for the linear model and the nonlinear model with loose monetary policy and with tight monetary policy based on the deviation of the real WIBOR from its own stochastic trend. We show results only for even horizons to conserve space. First, we check whether the instruments used in the LP-IV estimation are weak or not weak at a given horizon *h*. If the Kleibergen-Paap rk F-statistic, listed under "F-statistic" in Table 1, is below 10, we consider it a weak instrument case. Next, we test for the nonlinear model at every horizon whether the multiplier under the tight monetary policy regime is the same as under the loose monetary policy regime. The p-value for the test of the null hypothesis that the multipliers are the same across tight and loose regimes is given in the last column of Table 1. At every horizon, the tight monetary policy state estimates are subject to weak instruments and we hence report Anderson-Rubin (AR) p-values, which are robust to weak instruments. All p-values are above 0.14, so that the null hypothesis that the multipliers are equal across states cannot be rejected. This is consistent with the state-dependent cumulative impulse responses shown in Figure 1 in the second graph moving closely together. Therefore, the stance of monetary policy has no statistically significant effects on the size of the government spending multiplier.

[13] See also Rossi and Zubairy (2011) for a large economy model for the U.S.
[14] The real exchange rate enters the model statistically significantly and excluding it from the model affects results, leading to spurious nonlinear multipliers.

The linear model specification, stated in equation (4), is supported by the data. Figure 1 reveals for the linear case that the effects of a government spending shock on real output are significant at a 5% level for all horizons up to h=9.[15] Also, the rk F-statistics for the linear model are above a value of 10 up to, and including, horizon h=8, and therefore multiplier estimates should be reliable up to that point. The cumulative multiplier increases after the impact of the shock from 0.56 steadily to reach a peak after four quarters with a value of 1.11. It then slowly falls to 0.82 at h=9 and becomes statistically insignificant afterwards. At the peak value, a 1 złoty increase in government purchases leads to a cumulative increase in real GDP of 1.11 złoty one year after the increase.

4.2 Results with the Monetary Policy Stance Based on the Monetary Condition Index

Table 2 presents results when the indicator variable for the baseline model is chosen based on the MCI of Section 3.2.2. Again, we first check whether the instruments used in the LP-IV estimation are weak or not at a given horizon. The p-value in the last column of Table 2 for horizons up to h=4 are based on HAC-adjusted covariances and for longer horizons on the weak-instrument robust AR test. The null hypothesis that government spending multipliers are the same in states of tight and loose monetary policy cannot be rejected at all horizons at significance levels above 0.17. Therefore, using the MCI to define the regime as tight or loose monetary policy in a given quarter instead of the real WIBOR interest rate leads to the same conclusion: the government spending multiplier does not depend on the stance of monetary policy and the linear case prevails. For convenience of comparison, Table 2 repeats the linear results, which are, of course, the same as for the linear case presented in Table 1 (and the linear case in Figure 4). .

4.3 Results with the Monetary Policy Stance Based on the Monetary Policy Rule

Table 3 reports results when the indicator variable for tight and loose monetary policy is chosen based on the monetary policy rule with interest rate smoothing of Section 3.2.3. The p-values for the Anderson-Rubin AR test for the null hypothesis that multipliers are the same across monetary policy states take on values above 0.36. Hence, there is no empirical evidence for nonlinearities in the government spending multiplier when this third alternative definition of tight and loose monetary policy is applied to our model. The linear model prevails.

[15] Ramey (2011, footnote 11, p. 11) points out that there is "no formal justification" provided in the literature for using 68% confidence intervals instead of the usual 95% intervals. Furthermore, Ramey mentions that it is neither justified on theoretical grounds to use 68% intervals.

4.4 Results for the Multiplier and the State of the Economy: The Output Gap

If the government spending multiplier is not dependent on monetary policy it may nevertheless depend on the state of the economy. We use economic slack (the output gap) based on the HP filter as in Section 3.2.3 and identify 59 quarters with expansions and 33 quarters with slack.

Results are shown in Table 4. First, we assess again for each horizon whether the instruments are weak and choose the relevant HAC- or AR-based test. The p-values in the last column of Table 4 reveal that the null hypothesis that government spending multipliers are the same in periods of economic slack and expansion cannot be rejected at any horizon with all p-values above 0.22. Therefore, the government spending multiplier does not depend on the state of the economy and stays the same over the business cycle. This means that the linear model in Table 2 and Figure 4 is the empirically relevant specification for Poland over the business cycle. The government spending multiplier reaches a peak of 1.11 after one year, regardless of whether the economy experiences slack or expansion.[16]

Ramey and Zubairy (2018) study government spending multipliers for the U.S. economy over the business cycle. After a military news shock, the U.S. spending multiplier is almost identical across states, with estimates of 0.60 and 0.59 after two years, in times of high and low economic slack, respectively.[17] The difference across states is not statistically significant, i.e., the spending multiplier is not dependent on slack. Our results also indicate no difference in multipliers across states for Poland, however, our multiplier is above unity. In other words, our evidence for Poland suggests a true multiplier effect in the sense that a 1 złoty increase in government purchases leads to an increase in real GDP of more than 1 złoty.

5 Sensitivity Analysis

5.1 Separating Government Consumption Purchases from Government Investment

Total government expenditures used in the analysis so far are the sum of two components: government consumption purchases and government investment. Boehm (2020) estimates statistically significantly different multipliers for government investment (near zero) and government consumption

[16] Ramey and Zubairy (2018), among others, define economic slack based on the level of the unemployment rate relative to a threshold. We follow Barro and Redlick (2011) and use the sample median as the threshold, which is 9.65% in our case. We find that using this alternative measure of economic slack leads to the same conclusion: the data support the linear model specification. Detailed results are available from the authors on request. Furthermore, quarterly data for Polish private debt are not available so that we are unable to consider a nonlinear specification with private debt overhang as in Bernardini and Peersman (2018).

[17] On the other hand, when they use a Blanchard–Perotti shock the spending multiplier estimates are 0.68 and 0.30 in times of slack and expansion, 2 years after the shock, and the difference is statistically significant.

(approximately 0.8) in a panel of OECD countries. Hence, we test whether the government consumption purchases and government investment spending have the same cumulative effects on real GDP with Polish data. Table B.1 in Appendix B lists the p-values for the test for every horizon for the empirically relevant linear model.[18] The results unanimously support the hypothesis that the government consumption multiplier is the same as the government investment multiplier, using the usual 5% significance level.

5.2 An Alternative Stochastic Trend Specification: Hamilton's Filter

The HP filter that we used is a double-sided filter that looks backwards and forwards from a given time period. It is debatable whether this is desirable or not for the construction of the indicator variable (Bernardini and Peersman 2018). We use the values of the HP-filtered real WIBOR interest rate series to determine the stance of monetary policy, though we do not use the filtered series directly in the regressions. In addition, we use the HP filter to extract the stochastic trend from real GDP. Hamilton (2018) demonstrates that the HP filter can introduce spurious dynamic relationships that are not present in the underlying data generating process.[19] Hamilton proposes instead a filter based on a regression of the variable at date t on its four most recent values as of date $t - s$, where s is set equal to 8 for quarterly data, in order to filter out the cycle. We define the stochastic trend as the remainder after removing the cycle from the time series considered.

We select tight and loose monetary policy states based on the deviation of WIBOR from its stochastic trend and replace the HP trend of real GDP with Hamilton's when applying the Gordon and Krenn (2010) transformation. Otherwise, the model specification is identical to that for the baseline model with WIBOR in Section 4.1. Table B.2 and Figure B.1 present the results. The test for linearity of the government spending multiplier produces p-values equal to or above 0.20, except for h=5 with a value of 0.06. At a 5% significance level, the linear model cannot be rejected at each and every horizon. Figure B.1 shows that the accuracy of local projection estimates is reduced at longer horizons (Haug and Smith 2012; Li et al. 2022). The linear model with Hamilton's filter gives similar estimates for the cumulative multipliers to those in Table 1. The peak occurs 1½ years after the shock instead of after one year with the HP filter, and the peak values is somewhat larger, 1.18, instead of 1.11. The multipliers are no longer statistically significant after horizon h=7. The multiplier steadily increase after the shock from a value of 0.76 to reach its peak and then falls afterwards. Overall, we conclude

[18] The nonlinear model for each component of government spending is not supported by the data.
[19] See also the subsequent rebuttal by Hodrick (2020).

that our baseline results are robust to choosing a different filter: the cumulative government spending multiplier for Poland is linear and has a value above one, between 1.1 and 1.2.

5.3 Excluding the COVID-19 Period from the Sample

We include in our model specification a scaled dummy variable to account for the severity of the effects of COVID-19 on the Polish economy. It is possible that this variable does not capture the effects in a satisfactory way and the dynamics of COVID could be different from levels effects captured by the dummy. If that were the case, one would expect to see different multipliers in the pre-COVID period. The scaled dummy variable has non-zero values in the period 2020Q1-2021Q4. Hence, we exclude this period from our sample and run regressions for the period 1999Q1 to 2019Q4.

Table B.3 gives results without the COVID-19 period included in the sample. The nonlinear results show some slight variations when compared to Table 1 with the COVID-19 period and the stringency index included. However, the nonlinear specification is not supported by the data, as was the case in Table 1. The linear results, as reported, are almost all identical to those in Table 1. The peak value at horizon four is slightly larger at 1.12 (instead of 1.11 in Table 1). At horizon h=6, the value is 1.06, compared to 1.05 in Table 1. All other linear multiplies are the same.

5.4 The Role of Government Debt

As an additional sensitivity check we add the ratio of government debt to GDP as an extra variable to the set of control variables in equation (2) in the baseline model of Section 4.1, using the WIBOR interest rate for defining the stance of monetary policy. High debt levels relative to GDP may reduce the effectiveness of fiscal policy due to government credibility issues. However, Poland's constitution limits public debt to 60% of GDP, though calculations depend on exact definitions. For example, OECD figures for the year 2020 record Poland’s gross government debt as 77.4% of GDP and for 2019 as 63.5%.[20] Such a fiscal rule may restrict government spending sufficiently so that debt levels do not affect fiscal policy effectiveness. Indeed, formal tests in our baseline model with WIBOR at every horizon cannot reject the null hypothesis that the coefficient estimates on the lags of the debt-to-GDP ratio are all zero at conventional significance levels. In other words, the debt-to-GDP ratio plays no role at every horizon considered for the effects of government spending shocks on real GDP in Poland.

[20] See https://www.oecd.org/gov/gov-at-a-gl ance-2021-poland.pdf, last accessed 24 August 2022.

5.5 The Response of Private Investment to a Government Investment Shock

We replace real GDP as the dependent left-hand side variable in the linear model in equation (4), y_{t+h}, with real private-sector investment.[21] However, we include both real GDP and real private-sector investment among all the other lagged control variables used in the baseline model specification. We construct private-sector investment by taking total gross capital formation in the Polish economy and subtracting from it government sector investment.

Figure B.2 in Appendix B graphs the cumulative response of private investment to a 1 złoty government investment shock. The effects on private investment are statistically significant at the 5% level, with horizons h=3 to h=6 being borderline cases. The shock decreases private investment in the range from 0.10 złoty to 0.66 złoty. All multipliers are negative. This suggest that government investment crowds out private investment but it is only partial crowding out that occurs. Table 1 shows that government purchases of 1 złoty increase real GDP by 1.11 złoty, i.e., the multiplier is 1.11. In Section 5.1 the empirical results show that government consumption purchases have the same effects on real GDP as government investment. In other words, the multipliers are the same, so that a 1 złoty increase in government investment cumulatively increases GDP by 1.11 złoty after one year, even though private-sector investment cumulatively falls by 0.10 złoty over the same time period. This results is opposite of what Boehm (2020) finds in an OECD-country panel with the government investment multipliers near zero, implying nearly complete crowding out of private investment. However, the conventional view is instead that a higher stock of public capital is a complement to private investment and overall raises productivity and output (e.g., Bom and Ligthart 2014), and our findings are consistent with this view.

6 Conclusion

This paper studies the coordination of fiscal and monetary policy. To be precise, we explore whether the government spending multiplier depends on the stance of monetary policy. We use quarterly Polish data from 1999Q1 to 2021Q4. We employ local projections with instrumental variables (LP-IV) to directly estimate cumulative multipliers and impulse response functions for government spending shocks. In order to account for the effects of COVID-19 on the Polish economy, we include in the regressions a scaled dummy variable in the form of a Poland-specific composite COVID-19 stringency index.

[21] The nonlinear model in equation (2) is empirically not supported for the specification with private investment and government investment shocks.

We find the government spending multiplier in Poland is not statistically significantly affected by monetary policy being tight or loose. We find that a positive government spending shock of 1 złoty leads cumulatively to a peak increase in real output of 1.11 złoty after four quarters and then tapes off to 0.82 złoty nine quarters after the shock and becomes statistically insignificant subsequently. These results are robust to various alternative definitions of the stance of monetary policy and also to the state of the economy being slack or expansion. Furthermore, we find that government spending on consumption and investment purchases have the same effects on real output, i.e., their multipliers are 1.11 as well. This is in contrast to some other empirical studies that find that multipliers differ depending on the type of government spending (e.g., Boehm 2020). In addition we find that excluding the COVID-19 period or including the ratio of public debt to GDP produces almost identical multipliers.

The impact of government investment on private-sector investment reveals that the effects are negative over all 12 horizons, with the effect at the fourth quarter taking a value of -0.10. Hence, there is some crowding out of private investment but it is relatively small and the overall peak effect of a government investment shock, having the same multiplier as total government spending, is above unity, 1.11.

Consistent with Ramey and Zubairy (2018), among others, the government spending multiplier in Poland is not state-dependent, however, we estimate in comparison a relatively large multiplier of 1.11.[22] A potential limitation of our analysis is that government investment shocks in our model specification are one-time shocks, occurring in one quarter only. Government investment sustained over several quarters may produce different results. Future research could include capital stocks and economic growth in an empirical model to study the longer-run effects of government investment on real output.

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[22] Ramey and Zubairy (2018) estimate the U.S. government spending multiplier to be around 0.60 after a military news shock.

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Table 1. Monetary Policy Stance Based on Real WIBOR Interest Rates

Horizon (h)	Linear Model		State-Dependent Model				p-Value for the Hypothesis Test†
			Tight Monetary Policy		Loose Monetary Policy		
	Multiplier	F-statistic	Multiplier	F-statistic	Multiplier	F-statistic	
2	0.72**	86.73	0.19	3.52	1.23	115.02	0.25 (AR)
	(0.26)		(0.45)		(0.15)		
4	1.11**	49.00	1.05	0.87	1.30	107.39	0.83 (AR)
	(0.21)		(0.59)		(0.08)		
6	1.05**	22.95	1.45	0.54	1.05	87.85	0.83 (AR)
	(0.14)		(1.14)		(0.06)		
8	0.93**	11.92	1.70	0.15	0.89	24.16	0.75 (AR)
	(0.25)		(2.00)		(0.06)		
10	0.66*	7.79	1.15	0.37	0.62	21.79	0.67 (AR)
	(0.34)		(0.62)		(0.11)		
12	0.35	6.07	1.16	0.33	0.32	13.69	0.51 (AR)
	(0.52)		(0.50)		(0.09)		

Notes: Multipliers reported are cumulative multipliers. Standard errors given in parentheses and are Newey-West heteroskedasticity and autocorrelation adjusted. 'F-statistic' refers to the first stage Kleibergen-Paap rk F-statistic statistic Weak instruments are identified by an F-statistic below 10 (Staiger and Stock, 1997). †The null hypothesis is that the multipliers across states of loose and tight monetary policy are not significantly different from each other: AR indicates Anderson-Rubin p-values, which are reported when the rk F-statistic is below 10 for at least one of the states. For the empirically supported linear model, * denotes statistical significance at the 10% level and ** significance at the 5% level.

Table 2. Monetary Policy Stance Based on the Monetary Condition Index

Horizon (h)	Linear Model		State-Dependent Model				p-Value for the Hypothesis Test†
			Tight Monetary Policy		Loose Monetary Policy		
	Multiplier	F-statistic	Multiplier	F-statistic	Multiplier	F-statistic	
2	0.72**	86.73	0.65	12.41	1.18	73.58	0.17 (HAC)
	(0.26)		(0.19)		(0.45)		
4	1.11**	49.00	1.21	11.93	1.21	15.08	0.92 (HAC)
	(0.21)		(0.29)		(0.45)		
6	1.05**	22,95	1.44	5.87	1.08	12.83	0.59 (AR)
	(0.14)		(0.45)		(0.30)		
8	0.93**	11.92	1.27	1.24	0.76	5.58	0.42 (AR)
	(0.25)		(0.26)		(0.31)		
10	0.66*	7.79	1.52	0.25	0.27	4.85	0.34 (AR)
	(0.34)		(0.36)		(0.39)		
12	0.35	6.07	6.58	0.00	-0.12	5.20	0.92 (AR)
	(0.52)		(246.80)		(0.54)		

Notes: See Table 1. †HAC indicates Newey-West p-values.

Table 3. Monetary Policy Stance Based on the Monetary Policy Rule with Interest Rate Smoothing

Horizon (h)	Linear Model		State-Dependent Model				p-Value for the Hypothesis Test†
			Tight Monetary Policy		Loose Monetary Policy		
	Multiplier	F-statistic	Multiplier	F-statistic	Multiplier	F-statistic	
2	0.72**	86.73	1.22	11.11	1.93	9.77	0.40 (AR)
	(0.26)		(0.58)		(0.51)		
4	1.11**	49.00	1.96	0.97	2.26	6.65	0.89 (AR)
	(0.21)		(1.75)		(0.47)		
6	1.05**	22,95	1.38	0.25	1.24	5.11	0.95 (AR)
	(0.14)		(2.22)		(0.24)		
8	0.93**	11.92	0.01	0.11	0.99	1.58	0.72 (AR)
	(0.25)		(1.76)		(0.35)		
10	0.66*	7.79	-0.92	0.03	0.88	1.12	0.77 (AR)
	(0.34)		(5.41)		(0.24)		
12	0.35	6.07	-5.64	0.03	1.30	1.83	0.36 (AR)
	(0.52)		(20.09)		(0.37)		

Notes: See Table 1.

Table 4. Business Cycle States Based on the Output Gap

Horizon (h)	Linear Model		Business Cycle State-Dependent Model				p-Value for the Hypothesis Test†
			Economic Expansion		Economic Slack		
	Multiplier	F-statistic	Multiplier	F-statistic	Multiplier	F-statistic	
2	0.72**	86.73	1.35	33.66	1.09	15.63	0.56 (HAC)
	(0.26)		(0.48)		(0.13)		
4	1.11**	49.00	1.26	29.23	1.26	8.17	0.98 (AR)
	(0.21)		(0.59)		(0.34)		
6	1.05**	22,95	0.69	30.33	0.87	5.03	0.71 (AR)
	(0.14)		(0.40)		(0.20)		
8	0.93**	11.92	0.46	21.11	0.67	3.74	0.67 (AR)
	(0.25)		(0.26)		(0.16)		
10	0.66*	7.79	0.44	9.38	0.52	3.89	0.85 (AR)
	(0.34)		(0.11)		(0.16)		
12	0.35	6.07	0.60	6.81	0.17	5.20	0.39 (AR)
	(0.52)		(0.19)		(0.20)		

Notes: See Tables 1 and 2.

Figure 1. State of Monetary Policy as Identified by Deviations from the Stochastic Hodrick-Prescott Trend in Relation to Real WIBOR

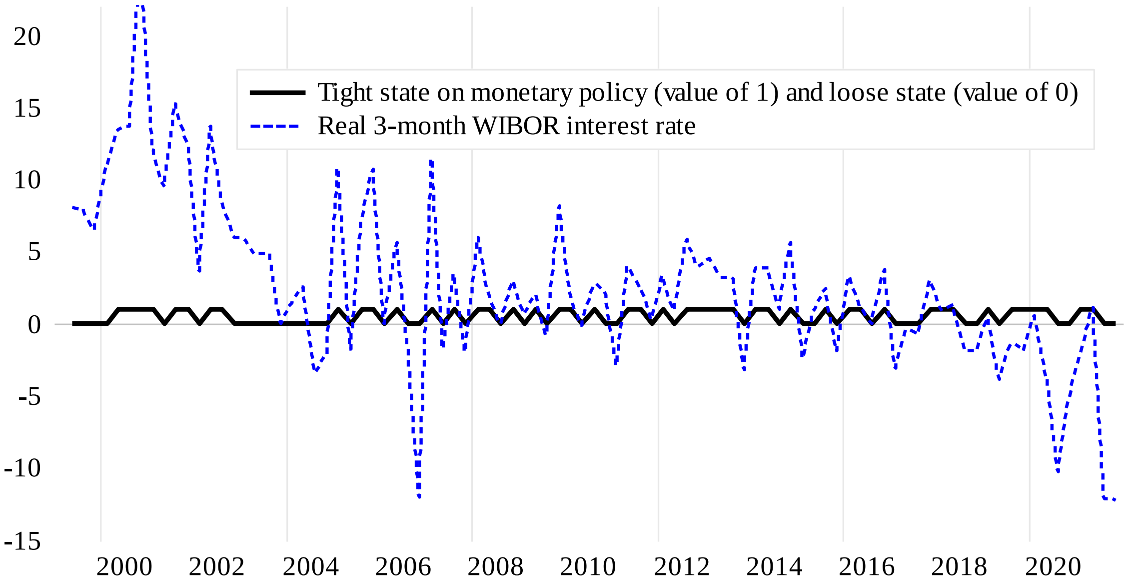


Figure 2. Monetary Condition Indices

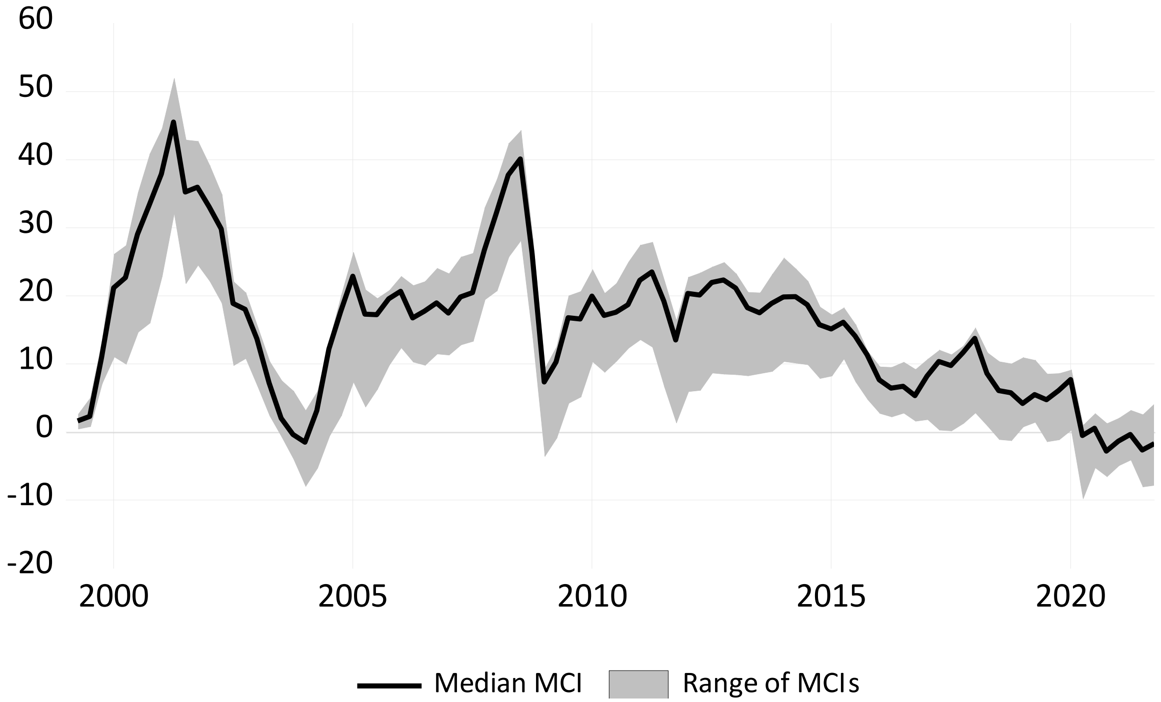
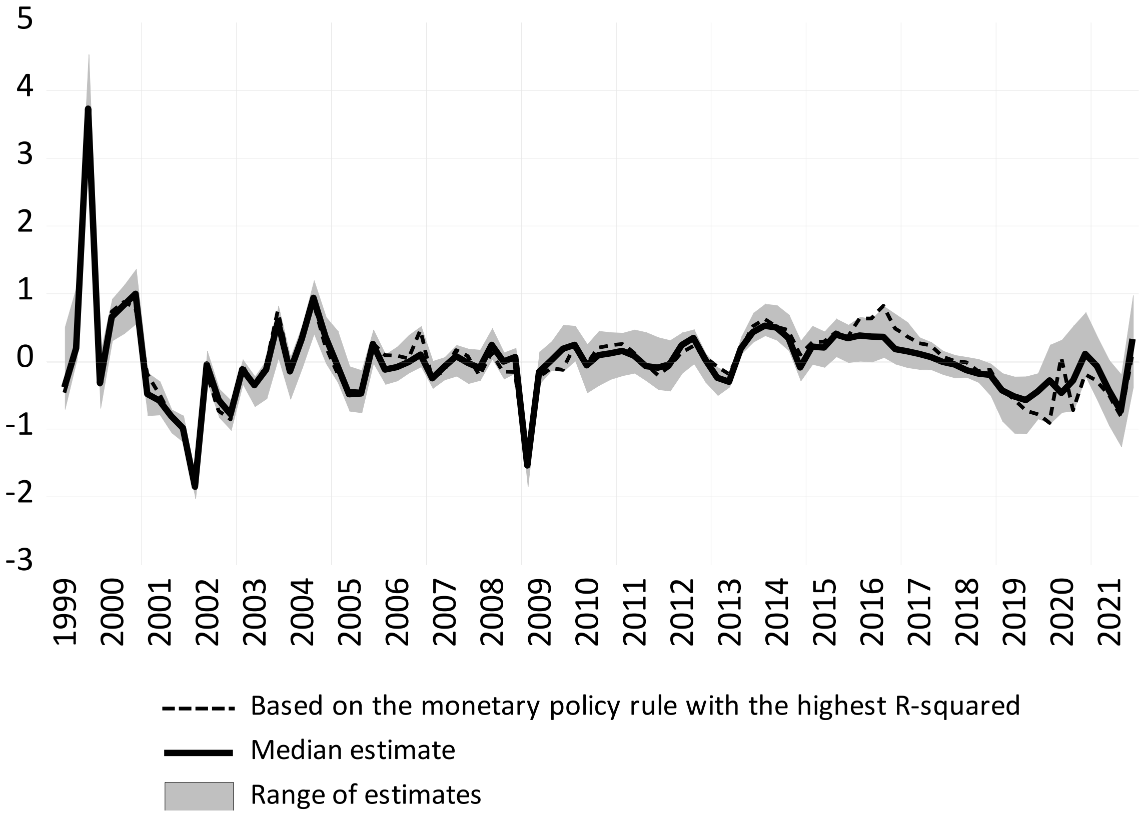
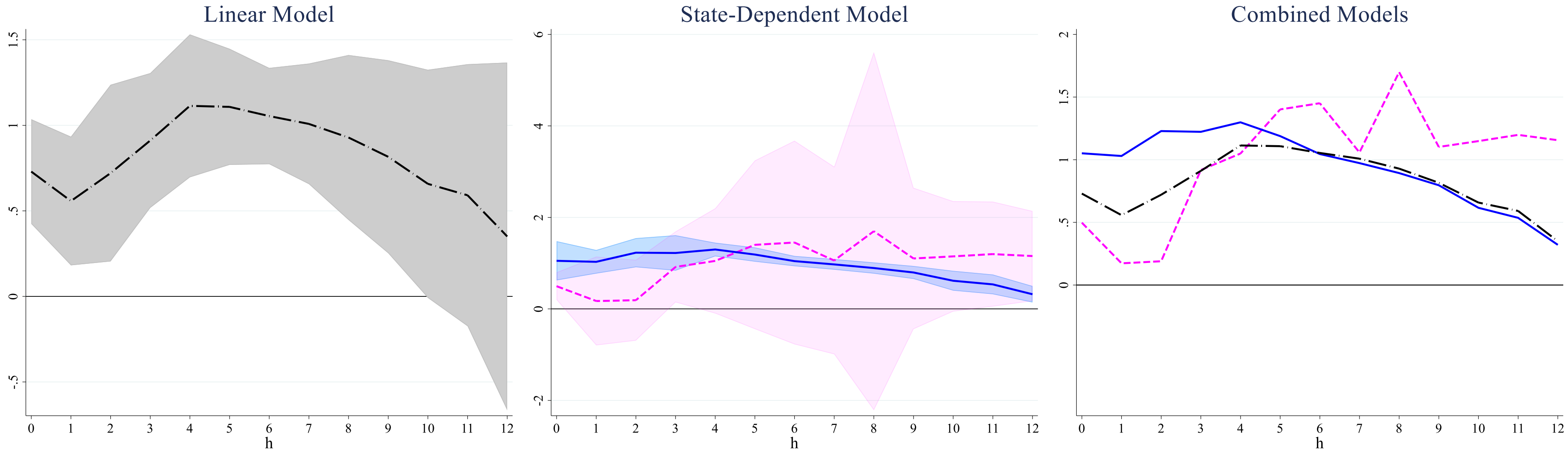


Figure 3. Interest Rate Gaps Based on Monetary Policy Reaction Functions with Interest Rate Smoothing



Notes: This figure presents differences between the short-term interest rate (3-month WIBOR) and theoretical interest rates derived on the basis of monetary policy rules with interest rate smoothing. Positive values correspond to periods of excessively restrictive monetary policy, given its typical response function, while negative values correspond to excessively loose monetary policy.

Figure 4. Responses of GDP (in złoty) to a Government Spending Shock When the Stance of Monetary Policy is Identified by the Deviation of the Real Interest Rate from its Stochastic Trend



Notes: This figure presents the cumlative impulse repsonses of real GDP for the linear and state-dependent models, where h refers to the horizon. The first two graphs are presented with shaded 95% confidence intervals based on Newey-West HAC standard errors. The third gaph displays the IRFs from each model for comparison. The blue (solid) line depicts the loose monetary policy state when the real interest rate is below its stochastic trend and the magenta (broken) line the tight monetary policy state when it is above its stochastic trend.

Appendix A: Data and Code

Data are from Eurostat (https://ec.europa.eu/eurostat/databrowser/explore/all/all_themes), unless otherwise indicated below. Data were retrieved on various days in April and May of 2022. Stata codes and the data are available from the authors on request. The abbreviation refer to the Stata programmes adapted by the authors from the Stata code posted by Ramey and Zubairy (2018) at https://www.journals.uchicago.edu/doi/suppl/10.1086/696277 (Data Archive; last accessed on 12 July 2022). Their codes *jordagk.do* and *jordagk_ar.do* were modified and used. The data definitions and sources are:

- rgdp - Real GDP, chain linked volumes (2010), million units of national currency (B1GQ from NAMQ_10_GDP);
- gdpd - GDP deflator, price index (implicit deflator), 2010=100, national currency (B1GQ from NAMQ_10_GDP) ;
- nom_ginv - Gross fixed capital formation of general government (P51G from gov_10q_ggnfa) – seasonally adjusted by the authors with the X12 option in EViews;
- nom_inv_total – Total gross fixed capital formation, seasonally adjusted series (P51G from NAMQ_10_GDP);
- nom_gcons - Final consumption expenditure of the general government (P3_S13 from gov_10q_ggnfa) – seasonally adjusted by the authors with the X12 option in EViews;
- nom_gov_tax_rev - Government tax revenue (D2REC+D5REC from gov_10q_ggnfa) – seasonally adjusted by the authors with the X12 option in EViews;
- nom_gov_transf - Government social transfers payable (D62_D632PAY from gov_10q_ggnfa) – seasonally adjusted by the authors with the X12 option in EViews;
- nom_gov_subs - Government subsidies payable (D3PAY from gov_10q_ggnfa) – seasonally adjusted by the authors with the X12 option in EViews;
- public_debt – Nominal government consolidated gross debt, (GD from gov_10q_ggdebt);
- wibor – WIBOR 3M, short-term 3-month interest rate relevant for Polish monetary policy (IRT_ST_M);
- ntwi - nominal effective exchange rate; Bank for International Settlements, BIS Statistics Warehouse (https://stats.bis.org/#);
- rtwi - real effective exchange rate; Bank for International Settlements, BIS Statistics Warehouse (https://stats.bis.org/#);

- si – COVID-19 stringency index. The stringency index is a composite measure based on nine response indicators including school closures, workplace closures, and travel bans, rescaled to a value from 0 to 1 (1 = strictest); Our World in Data (https://ourworldindata.org/);
- unemp – Unemployment rate, percentage of population in labour force, 15 to 74 years age class, seasonally adjusted series (UNE_RT_Q);
- monetary policy stance dummies – own calculations based on data from Eurostat and GUS (Central Statistical Office, Poland).

Appendix B: Sensitivity Analysis

Table B.1. P-Values for Testing of the Null Hypothesis of Equivalence of Multipliers for Government Consumption and Government Investment Purchases

Horizon (h)	Linear Model
1	0.09
2	0.21
3	0.24
4	0.61
5	0.74
6	0.88
7	0.90
8	0.75
9	0.37
10	0.16
11	0.10
12	0.11

Notes: The table reports p-values for a Wald test in the linear model that the cumulative multiplier for a government consumption shock is the same as the multiplier for a government investment shock.

Table B.2. Monetary Policy Stance Based on WIBOR: Hamilton's (2018) Filter

Horizon (h)	Linear Model		State-Dependent Model				p-Value for the Hypothesis Test†
			Tight Monetary Policy		Loose Monetary Policy		
	Multiplier	F-statistic	Multiplier	F-statistic	Multiplier	F-statistic	
2	0.92**	30.60	0.30	23.42	0.65	19.04	0.56 (HAC)
	(0.33)		(0.14)		(0.57)		
4	1.12**	21.12	0.05	7.93	0.82	5.77	0.20 (AR)
	(0.30)		(0.14)		(0.46)		
6	1.18**	26.82	0.11	4.69	1.08	5.38	0.27 (AR)
	(0.27)		(0.21)		(0.40)		
8	0.56*	17.83	-0.01	1.20	0.41	2.69	0.73 (AR)
	(0.43)		(0.44)		(0.44)		
10	-0.80	11.70	-1.02	0.75	-0.78	1.99	0.89 (AR)
	(1.10)		(1.29)		(1.13)		
12	-1.95	5.86	-2.54	0.42	-3.51	0.64	0.77 (AR)
	(1.86)		(2.93)		(3.10)		

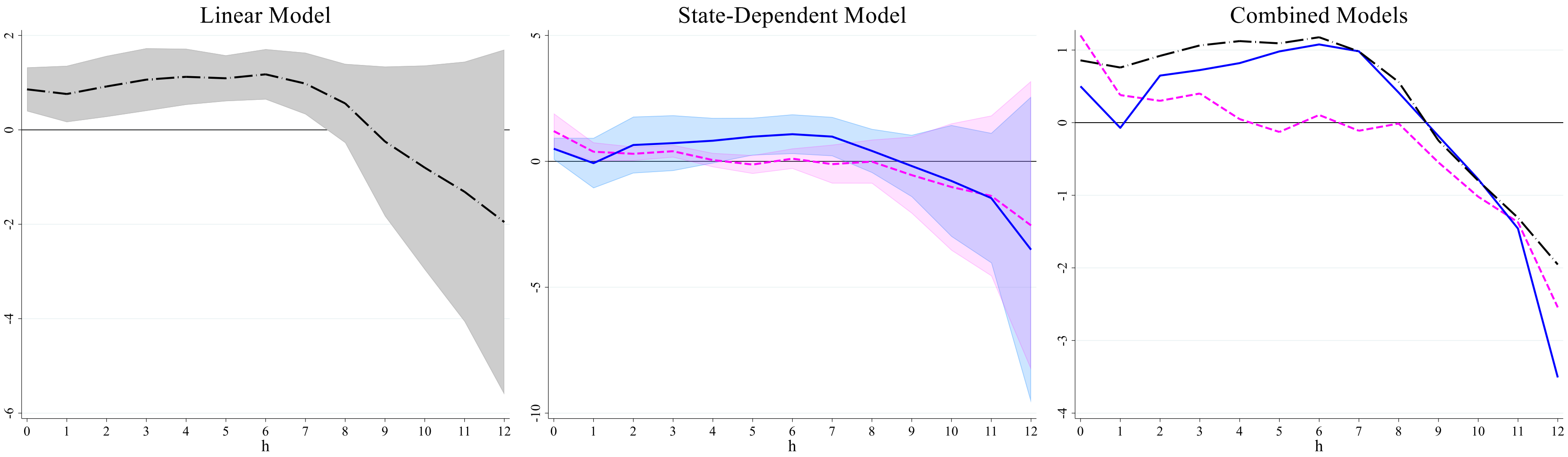
Notes: See Tables 1 and 2. Hamilton's (2018) filter was applied to extract the stochastic trend from real GDP and also from the real WIBOR interest rate instead of the Hodrick-Prescott filter, which was used in Table 1 for both variables.

Table B.3. Monetary Policy Stance Based on WIBOR: Excluding the COVID-19 Period

Horizon (h)	Linear Model		State-Dependent Model				p-Value for the Hypothesis Test†
			Tight Monetary Policy		Loose Monetary Policy		
	Multiplier	F-statistic	Multiplier	F-statistic	Multiplier	F-statistic	
2	0.72**	92.87	0.38	4.63	1.23	102.92	0.25 (AR)
	(0.26)		(0.50)		(0.15)		
4	1.12**	59.26	1.53	0.93	1.25	109.62	0.83 (AR)
	(0.21)		(0.83)		(0.07)		
6	1.06**	26.21	1.86	0.60	1.05	90.88	0.83 (AR)
	(0.16)		(1.37)		(0.06)		
8	0.93**	11.92	1.70	0.15	0.89	24.16	0.75 (AR)
	(0.25)		(2.00)		(0.06)		
10	0.66*	7.79	1.15	0.37	0.62	21.79	0.67 (AR)
	(0.34)		(0.62)		(0.11)		
12	0.35	6.07	1.16	0.33	0.32	13.69	0.51 (AR)
	(0.52)		(0.50)		(0.09)		

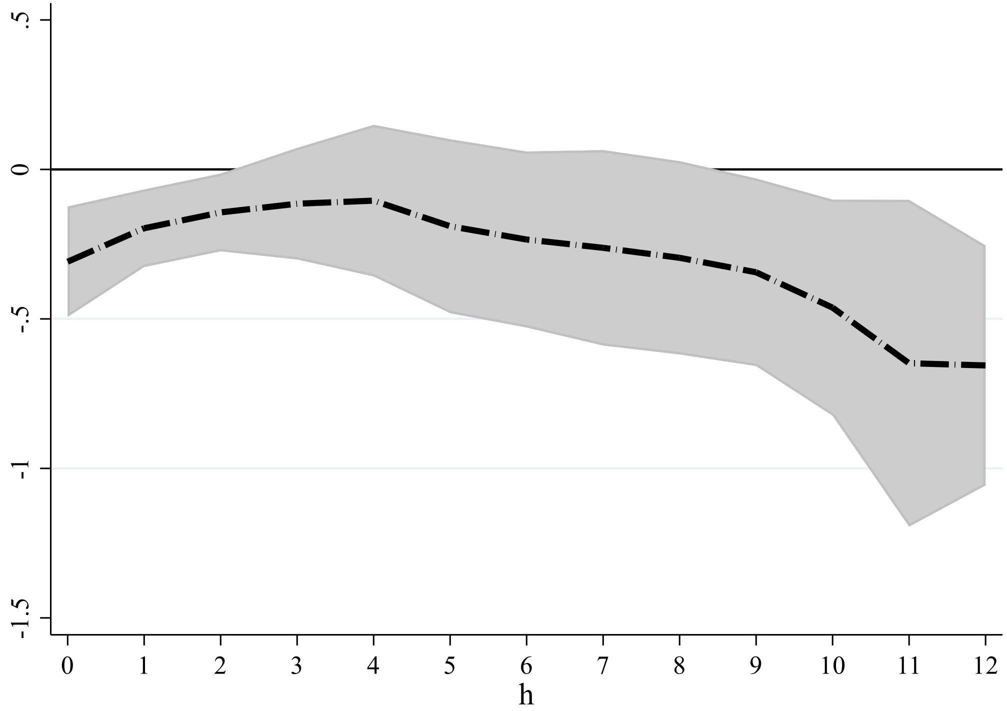
Notes: See Table 1.

Figure B.1. Responses of GDP (in złoty) to a Government Spending Shock When the Stance of Monetary Policy is Identified by the Deviation of the Real Interest Rate from its Stochastic Trend: Hamilton's (2018) Filter



Notes: See Figure 4 and Table B.2.

Figure B.2. The Response of Private Investment to a Government Investment Shock



Notes: See Figure 4.