Nonlinearities and state-dependence in the monetary transmission mechanism: Evidence from a commodity-dependent economy

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Abstract

This paper examines nonlinearity and state-dependence in the transmission mechanism of monetary policy in Mongolia, a commodity-dependent developing economy, using local projection methods. Our empirical work resulted in several novel findings. First, transmission lags vary with economic conditions, with longer lags during recessions compared to expansions. Second, the effectiveness of monetary policy in stimulating GDP is enhanced during economic recessions, the loosening phase, and high inflation regimes, while its ability to control inflation is diminished in expansions. Monetary policy pass-through to bank interest rates is more pronounced during the joint state of recessions and monetary tightening. Third, expansionary monetary policy shocks have a larger impact, leading to a depreciation of the real effective exchange rate and sharp declines in bank interest rates. These findings are robust across various model specifications, highlighting the importance of economic states for the effectiveness of monetary policy.

Keywords: Monetary policy transmission, Interest rate pass-through, State dependence, Business cycle, Inflation regimes, Local projections

JEL classifications: E32, E31, E32, E43, E52, C32.

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1. Introduction

The New Keynesian framework serves as a key foundation for analyzing business cycle fluctuations and the effects of monetary policy. A core feature of this framework is the empirically supported assumption of price stickiness, which explains why monetary policy can influence real economic variables. Monetary policy decisions affect economic activity and inflation through various channels, collectively referred to as the monetary transmission mechanism, which is characterized by long, variable, and uncertain time lags. These transmission channels have been extensively discussed in the literature (e.g., Mishkin 1996, Bernanke and Gertler 1995, Beyer et al. 2024). Recent empirical studies (i.e., Ascari and Haber 2022 for U.S. data, Alpanda et al. 2021 for 18 advanced economies, and Berger et al. 2021 and Eichenbaum et al. 2022 for U.S. data) highlight that the effects of monetary policy are state-dependent. However, the existing literature primarily focuses on advanced economies with closed economy settings and has not explored the state-dependent transmission of monetary policy in open economies, particularly in the context of emerging markets and developing economies (EMDEs).

In this regard, our paper empirically investigates the state dependence and asymmetry in the transmission mechanism of monetary policy in Mongolia, one of the most commodity-dependent EMDEs. Specifically, we examine i) how different states of the business cycle, inflation, and policy rate influence the monetary policy transmission (i.e., economic activity, inflation, exchange rate, and interest rate pass-through), and ii) how positive and negative changes in the policy rate asymmetrically affect the transmission mechanism using the local projection approach proposed by Jordá (2005). This methodology is particularly well-suited for analyzing state-dependent effects, as it allows the parameters to change according to the state of the economy. The empirical analysis focuses on Mongolia, a small, open, and commodity-exporting economy that encounters significant challenges in managing the effects of external shocks and improving the effectiveness and framework of its monetary policy. Therefore, the findings from this case study can be extended to other commodity-dependent EMDEs that face similar structural and macroeconomic challenges.

This paper makes two key contributions to the literature. First, it is one of the earliest attempts to examine the state-dependent effects of monetary policy across business, inflation, and interest rate cycles using local projection methods, focusing on a commodity-dependent developing economy. This approach enables a comprehensive analysis of how monetary policy influences exchange rate dynamics and interest rate pass-through under different economic states. Second, by documenting the state-dependent and asymmetric effects of monetary policy shocks, the paper provides evidence-based policy insights aimed at enhancing the effectiveness of monetary policy in commodity-dependent EMDEs.

The existing literature provides theoretical and empirical foundations for analyzing the state-dependent macroeconomic effects of monetary policy shocks across business, inflation, and interest rate cycles within the unified framework as employed in this paper. Several papers (i.e., Santoro et al. 2014, Bernstein 2021, Alpanda et al. 2021, Eichenbaum et al. 2022, Kloosterman et

al. 2024) have developed theoretical foundations and dynamic general equilibrium models in explaining state-dependences of monetary policy across varying economic conditions.

Recent empirical studies have explored the state-dependent effects of monetary policy in advanced economies using local projection methods[2]. Extensive empirical evidence indicates that output and prices respond asymmetrically to monetary policy shocks across different phases of the business cycle. However, the empirical findings remain ambiguous. Santoro et al. (2014) find that monetary policy has stronger effects on GDP during contractions, as compared with expansions, and there is no difference in price responses across different stages of the cycle. In contrast, Tenreyro and Thwaites (2016) demonstrate that (i) monetary policy is less effective during recessions, and (ii) contractionary policy shocks have a greater impact than expansionary shocks. Using a larger sample on an annual basis for 17 advanced economies, Jordà et al. (2020) document a weaker output response to monetary policy shocks during slumps compared to booms, while the evidence of asymmetry is less clear for the inflation response.

State-dependent responses of macroeconomic and financial variables to monetary policy shocks across different phases of inflation and financial cycles have gained increased attention in recent research. Relying on the sticky price theory of monetary policy transmission, Ascari and Haber (2022) find that the effects of monetary policy shocks on industrial production, inflation, and the federal funds rate depend on both shock size and trend inflation. Their results indicate that inflation responds more strongly after a large shock and during high-trend inflation regimes, supporting the argument that state-dependent pricing is a vital feature of the transmission mechanism of monetary policy. Using a panel data set from 18 advanced economies, Alpanda et al. (2021) find that the impacts of monetary policy shocks on macroeconomic and financial variables are weaker during periods of economic downturns, low household debt, and high interest rates. Berger et al. (2021) argue that the effectiveness of stimulating the economy through interest rate cuts in the presence of substantial debt in fixed-rate, prepayable mortgages depends not only on the current level of interest rates but also on their prior previous trajectories. Jordà et al. (2020) show that (i) the responses of real GDP and the price level to monetary policy shocks are notably stronger when inflation exceeds 2 percent, and (ii) household leverage (i.e., mortgage credit cycle) influences the effectiveness of monetary policy in ways that lending to firms does not. In contrast to the existing literature on the state-dependent interest rate pass-through using local projections, we consider the state dependence of interest rate pass-through along all three cycles (i.e., business, inflation, and policy rate cycles). For instance, Puglisi (2015) presents empirical evidence for state-dependent pass-through from the policy rate to lending rates in the US. He argues that a high initial skewness

[2] Local projection methods are well reviewed and discussed by several papers (i.e., Jordà 2005, 2023, Plagborg-Møller and Wolf 2021, Olea and Plagborg-Mølle 2021, and Li et al. 2024). Other methods such as a logit mixture vector autoregressive (Burgard et al. 2019) and single regressions with product variables (Eichenbaum et al. 2022) have also been used in analyzing the state-dependent transmission of monetary policy. Burgard et al. (2019) find that both output and prices are found to decrease after monetary policy shocks, and the contraction is much stronger especially during "crisis times" as the peak effect is roughly one-and-a-half times as large when compared to "normal times". Eichenbaum et al. (2022) reveal that the efficacy of monetary policy is state-dependent, varying in a systematic way with the pool of potential savings from refinancing.

in the cross-sectional distribution of lending rates across banks (prior to a change in the policy rate) results in a stronger response of bank lending rates and economic activity. Messer and Niepmann (2023) find that pass-through of policy rates to deposit rates in the euro area during the current tightening cycle (July 2022-March 2023) has been more sluggish than in previous tightening episodes and show that levels of excess reserves explain cross-country variation in the pass-through. Using data on seven different bank interest rates in 30 European countries, Beyer et al. (2024) show that monetary policy pass-through to bank deposit and lending rates in the post-pandemic hiking cycle has been heterogenous across countries and types of interest rates and has generally been weaker and slower.

Though several papers have examined the transmission mechanism of monetary policy in Mongolia, no paper has employed local projection methods to analyze the state-dependent effects of monetary policy shocks. Using quarterly vector autoregressions (VARs), Bayardavaa et al. (2015) find that i) a change in the interbank market rate significantly affects bank lending rate with time lags of 1-2 quarters, and ii) bank lending is the most effective channel of monetary transmission to output and prices. Employing both linear and nonlinear monthly autoregressive distributed lag (ARDL) models, Gan-Ochir (2016) demonstrates that (i) the long-run monetary policy pass-through to bank deposit rates is stronger but slower compared to the pass-through to bank lending rates, and (ii) the pass-through to deposit and lending rates is asymmetric, with lending rates responding more strongly to policy rate cuts, while deposit rates show a stronger response to policy rate hikes. Using sign-restricted VARs, Gan-Ochir and Davaasukh (2023) find that (i) the transmission lag to the peak effect on the consumer price index (CPI) is 11-12 months, with the maximum impact reaching 1 percent following a 1.1 percentage point shock to the policy rate, and (ii) bank deposits and newly issued loans serve as key transmission channels of the monetary policy. Based on Bayesian VARs, Gan-Ochir (2023) also observes hump-shaped responses of GDP and CPI to a policy rate shock, along with a strong response of bank loans to the shock. Gan-Ochir and Munkhbayar (2024) demonstrate that bank liquidity creation and its components respond significantly to a policy rate shock, with peak effects occurring after time lags of 3-4 quarters. Using Bayesian estimated structural models, several studies (e.g., Gan-Ochir and Undral 2018, and Gan-Ochir and Munkhbayar 2023) have analyzed the transmission channels and effects of policy rate shocks on macroeconomic and financial variables. These studies reveal that the policy rate pass-through to bank lending rates is approximately 20 percent, with a lag of 2-3 quarters, and that the maximum effects on real wages, consumption, investment, output, and inflation occur with lags of 3-4 quarters.

The remainder of the paper is organized as follows. Section 2 provides an overview of the Mongolian economy and the functioning of monetary policy. Section 3 outlines the empirical methodologies based on local projections. Section 4 describes the data, presents the main findings, and conducts robustness checks. Finally, Section 5 concludes the paper and explores policy implications.

2. Overview of the Mongolian economy and the operation of monetary policy

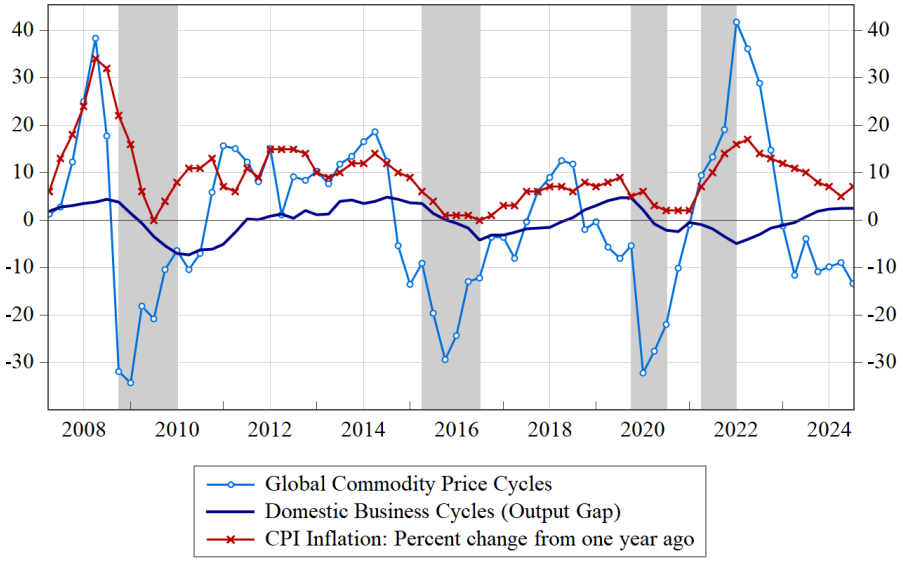
The Mongolian economy is heavily dependent on its mineral exports, making it highly vulnerable to external shocks such as fluctuations in foreign demand for resources, global commodity prices, and foreign direct investment (FDI). Commodity exports, including minerals and animal products, account for approximately 95 percent of the country's total exports, equivalent to 40 percent of its gross domestic product (GDP)[3]. People's Republic of China is a big trading partner of Mongolia as the trade between the two countries accounts for 90 percent of total exports and 40 percent of imports. Imports from Russia account for about 30 percent of total imports. Historically, FDI inflows have been predominantly directed toward the mining sector. Due to its lack of economic diversification and susceptibility to boom-and-bust cycles, fluctuations in export prices and volumes of commodities have caused significant macroeconomic volatility in the economy.

The economic profile of Mongolia exhibits several typical characteristics shared by commodity-dependent emerging and developing economies. The import-to-GDP ratio is high, at approximately 65 percent, with nearly 50 percent of the household consumption basket comprising imported goods. Mongolia's trade landscape is marked by a significant reliance on the U.S. dollar, referred to as "dominant currency pricing". Notably, the US dollar accounts for a substantial share (around 70 percent) of critical economic metrics, including imports and external debt payments. In 2023, Mongolia's external debt was equivalent to approximately 170 percent of its GDP, a level significantly higher than those observed in low-income developing countries in the Middle East and North Africa (MENA) region, as well as in emerging markets and middle-income countries in the Caucasus and Central Asia (CCA). Mongolia also shares several economic characteristics with these regions, including relatively high dollarization (25 percent for bank deposits), shallow foreign exchange markets, imperfect credit markets, underdeveloped stock markets, balance sheet currency mismatches in both private and public sectors, reliance on bank credits (bank credits to GDP ratio is about 40 percent), poorly anchored inflation expectations, immature monetary policy frameworks, and the imperative to enhance credibility in the economy (Gan-Ochir et al. 2024).

Mongolia's business cycles have been primarily driven by global commodity price fluctuations. Over the last two decades, the economy has experienced four recessions. The first three—during the Global Financial Crisis (GFC), the 2014–2015 downturn, and the onset of the COVID-19 pandemic—were directly linked to declines in global commodity price cycles. However, the most recent recession, which occurred during 2021-2022, was driven by a decline in coal export volumes due to China's "Zero-COVID" policy, rather than changes in commodity prices (Figure 1).

[3] The percentage of natural resource exports in the economy is in line with that of most low-and lower-middle-income resource-rich countries (where the share of exports from natural resources is above 70 percent).

Figure 1. Global Commodity Price Cycles, Domestic Business Cycles, and CPI Inflation



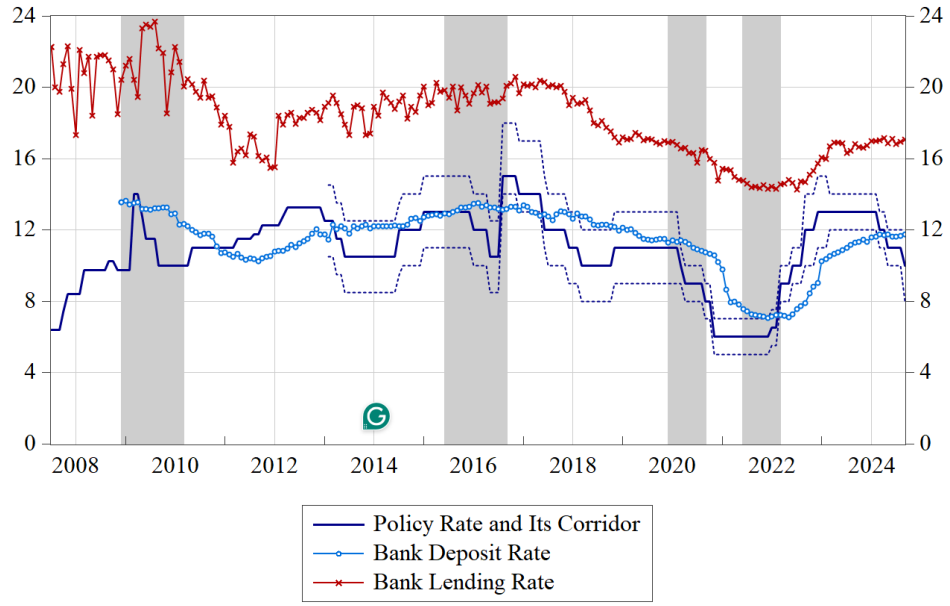
Notes: Commodity price cycles and domestic business cycles (GDP gap) are calculated in percentage points using the Hodrick-Prescott (HP) filter for quarterly data. The shaded area represents the recession period identified by the Bry-Boschan (BBQ) business cycle dating algorithm.

Source: Authors' calculation using the data retrieved from the National Statistical Office of Mongolia and IMF Primary Commodity Prices.

Inflation in Mongolia has been volatile and relatively high over the past two decades. Its dynamics have generally mirrored domestic business cycles, indicating that demand factors play a significant role. Except for the 2021-2022 recession, inflation typically declines during recessions and rises during periods of economic expansion (Figure 1). The stagflation observed during the recent recession was largely driven by supply-side factors, including increased transportation costs due to the pandemic and higher world food and oil prices resulting from the Russia-Ukraine conflict. These negative supply shocks impacted the economy at a time when both monetary and fiscal policies were highly accommodative, aimed at stimulating economic activity during the pandemic. In addition to exchange rate fluctuations and pro-cyclical fiscal policy, which amplify external shocks, supply-side factors have been key drivers of inflation in Mongolia. These factors include changes in food and gasoline prices, high seasonality in the domestic supply of meat, wheat, and vegetables, supply bottlenecks at the borders, elevated logistical costs, wage growth, and severe weather conditions such as dzud (harsh winters) and droughts (Gan-Ochir and Davaajargal 2023)[4].

[4] They find that external variables (i.e., China's growth, China's inflation, change in oil price) play an important role in forecasting inflation, and among domestic variables, wage inflation, and M2 growth are the best predictors for short and longer forecasting horizons.

Figure 2. Interest rates in Mongolia, in percentage points



Notes: The shaded area represents the economic recession period.
Source: The Bank of Mongolia's Monthly Statistical Bulletin.

The Central Bank of Mongolia (BOM) has been implementing a forward-looking inflation-targeting framework since 2007. As its main policy instrument, the BOM uses the policy rate[5] to signal its monetary policy stance[6]. In July 2007, the BOM introduced the policy rate as the new operational target for monetary policy, representing the desired level for the money market rate. The BOM's monetary policy operations focus on maintaining conditions in the money market to ensure the market rate remains close to the operational target (i.e., the policy rate). The policy rate is periodically adjusted by the Monetary Policy Committee (MPC). In February 2013, the MPC introduced an interest rate corridor system to strengthen the interest rate channel of monetary policy transmission. This system has made monetary policy significantly more transparent and easier to communicate, as it operates directly through the short-term interest rate. While Mongolia officially maintains a floating exchange rate regime, the BOM retains the authority to intervene in foreign exchange markets to reduce excessive volatility in the Mongolian Tögrög (MNT) exchange rate. Banks play a crucial role in the transmission mechanism of monetary policy, as the banking sector accounts for approximately 90 percent of the financial system's total assets (Gan-Ochir 2016).

As shown in Figure 2, when the monetary policy rate was introduced in mid-2007, the bank lending rate was notably high, at approximately 22 percent. This was primarily due to a combination of a

[5] In current practice, the policy rate is equal to the interest rate paid on the central bank bill with the shortest maturity (i.e, one work or four-weeks).
[6] In response to the economic recessions (i.e., the period 2013-2016 and the period 2020-2023), the BOM has also actively implemented unconventional monetary policy measures, classified by international financial institutions as quasi-fiscal policy activities (QFPA). The characteristics of QFPAs and their macroeconomic effects are discussed by Gan-Ochir and Davaasukh (2023).

high bank deposit rate and a significant intermediation spread, accounting for factors such as a high volume of nonperforming loans, elevated reserve requirement ratios, substantial market risks, inefficiencies in bank operations, and the low performance of financial intermediation. Since the policy rate has been set with the objective of maintaining macroeconomic stability, it tends to be maintained at lower levels during periods of economic recessions.

The general co-movements between the policy rate and bank interest rates (deposit and lending rates) have slightly improved over time. The establishment of the interest rate corridor system in 2013 has been instrumental in keeping money market rates within desired levels and enhancing the interest rates pass-through (Gan-Ochir 2016). However, changes in the policy rate have been relatively smooth compared to the volatilities observed in inflation and the output gap. In the empirical analysis presented in the subsequent sections, we formally investigate the interest rate pass-through and the effectiveness of monetary policy in stabilizing the macroeconomy under varying economic conditions.

3. Empirical methodology: Local projection approach

We consider both linear and non-linear local projection methods. All empirical models include 8 variables including year-on-year (YoY) growth in China's steel production (Δcsp_t), YoY growth in global commodity prices ($\Delta comp_t$), YoY GDP growth (g_t), YoY growth in real effective exchange rate (REER, $\Delta reer_t$), YoY CPI inflation (π_t), policy rate (pr_t), deposit rate (dr_t), and lending rate (lr_t). The inclusion of the first two variables is motivated by Gan-Ochir and Davaajargal (2019) and Gan-Ochir et al. (2023) who show external shocks, particularly shocks to China's economy (demand for minerals) and global commodity prices, play an important role in Mongolia's business cycle fluctuations. The next four variables are standard variables included in a small open economy structural VAR models (i.e., Bjørnland 2008 for Norway, Dungey et al. 2020 for Australia, and Gan-Ochir et al. 2024 for Mongolia). In practice, policy rate decisions consider YoY inflation, GDP growth, and change in the exchange rate, and hence, we include the twelve-month log change in output, CPI, and REER. The last two variables are included in the models to explicitly examine state-dependent interest rate pass-through under varying economic conditions. The inclusion of bank rates is further motivated by VAR models used to study interest rate pass-through, such as those by De Bondt (2005) and Hristov et al. (2014) for the Euro area, and Gan-Ochir (2023) for Mongolia.

3.1 Linear local projections

In this section, we focus on a simple linear local projection (LP) in which IRFs can be defined as the difference between two forecasts:

$$IR(t, s, \mathbf{d}_i) = E(\mathbf{y}_{t+s}|\mathbf{v}_t = \mathbf{d}_i; \mathbf{X}_t) - E\left(\mathbf{y}_{t+s}\middle|\mathbf{v}_t = \underline{\mathbf{0}}; \mathbf{X}_t\right) \quad s = 0,1,2, \ldots \tag{1}$$

where the operator $E(.|.)$ denotes the best, mean squared error predictor; $\mathbf{y}_t$ is an $n \times 1$ random vector; $\mathbf{X}_t \equiv (\mathbf{y}_{t-1}, \mathbf{y}_{t-2}, \ldots)'$; $\underline{\mathbf{0}}$ is of dimension $n \times 1$; $\mathbf{v}_t$ is the $n \times 1$ vector of reduced-from

disturbances; and $\mathbf{D}$ is an $n \times n$ matrix, of which columns $\mathbf{d}_i$ are $n \times 1$ column vectors that contain the mapping from the structural shock for the i^{th} element of $\mathbf{y}_t$ to the experimental shocks (i.e., reduced form shocks).

To generate the impulse response functions (IRFs), we use the LP method proposed by Jordá (2005), which has certain advantages over the common vector autoregression (VAR). The LP technique is agnostic about the true DGP[7] and remains valid even when its Wold decomposition does not exist. It only requires projecting one period at a time, rather than an increasingly distant horizon as in the VAR[8].

The LP method calculates IRFs by running a sequence of forecast equations given by

$$\mathbf{y}_{t+s} = \boldsymbol{\alpha}^s + \mathbf{B}_1^{s+1}\mathbf{y}_{t-1} + \mathbf{B}_2^{s+1}\mathbf{y}_{t-2} + \cdots + \mathbf{B}_p^{s+1}\mathbf{y}_{t-p} + \mathbf{u}_{t+s}^s \quad s = 0,1,2,\dots,h \tag{2}$$

where $\mathbf{y}_t = [\Delta csp_t, \Delta comp_t, g_t, \pi_t, pr_t, \Delta reer_t, dr_t, lr_t]'$ is a vector of the model variables which we wish to forecast s step ahead (for h different forecast horizon), $\boldsymbol{\alpha}^s$ is a 8×1 vector of constants, $\mathbf{B}_i^{s+1}$ are 8×8 square matrices of coefficients for each lag i and horizon $s+1$, and $\mathbf{u}_{t+s}^s$ is the error term. A forecasting model consists of only p lags of the variables in the system[9].

According to definition (1), the IRFs from the local-linear projections in (2) are

$$\widehat{IR}(t, s, \mathbf{d}_i) = \widehat{\mathbf{B}}_1^s \mathbf{d}_i \quad s = 0,1,2,\dots,h \tag{3}$$

with the normalization $\mathbf{B}_1^0 = \mathbf{I}$, and $\widehat{\mathbf{B}}_1^s$ are the impulse response coefficients. Since the DGP is unknown, the valid inference for LP impulse responses can be obtained with heteroskedasticity and autocorrelation (HAC) robust standard errors. Hence, for sequential estimation, Jordá (2005) recommends using a Newey-West estimator (i.e., the most popular HAC robust estimator) of residual covariance matrix at each horizon step *s* to improve the accuracy of the confidence interval estimate. Alternately, Jordá (2009) proposes a joint estimation, in which this issue is parametrically addressed, and the correct covariance matrix of the full path of the impulse response coefficients is derived. To construct the confidence intervals for the IRFs, we employ the conditional error bands proposed by Jordá (2009). This approach is particularly effective in identifying the significance of individual impulse response coefficients, analogous to how a Wald test assesses the joint significance of coefficients.

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[7] Jordá (2005) presents that the IRFs generated by the LPs are equivalent to the ones calculated from a VAR when true DGP is a VAR, but the IRFs for other DGPs that are not true VARs, are better estimated using the LP method.

[8] The primary advantage over the standard VAR approach is its lack of structure from one horizon to the next. In the VAR approach, the IRFs at all horizons are directly connected to the VAR parameters. On the other hand, the LP method computes the IRFs from a different forecast equation, and thus the structure of the IRFs can vary over the horizon. This allows flexibility when the DGP is nonlinear (Ahmed and Cassou 2016).

[9] The maximum lag p can be determined by information criteria and need not to be common to each horizon s. The lag length and the dimension of the vector $\mathbf{y}_t$ will impose the degree of freedom constraints on the maximum practical horizon *h* for very small samples. Moreover, consistency does not require the sequence of *h* system regressions in (2) be estimated jointly-the IRFs for the *j*th variable in $\mathbf{y}_t$ can be estimated by a univariate regression of y_{jt} onto $\mathbf{X}_t \equiv (\mathbf{y}_{t-1}, \mathbf{y}_{t-2}, \dots, \mathbf{y}_{t-p})'$.

For constructing the experimental shock ($\mathbf{d}_i$), we use methods suggested by Jordá (2005), which follow techniques used in the traditional VAR literature[10]. This method starts by estimating a linear VAR and applying a Cholesky decomposition to the variance-covariance matrix. The Wold-causal order for the elements of $\mathbf{y}_t$ helps organize the triangular factorization of the reduced-form, residual variance-covariance matrix, $\mathbf{\Omega} = \mathbf{P}\mathbf{P}'$. Such an identification mechanism is equivalent to defining the experimental matrix as $\mathbf{D} = \mathbf{P}^{-1}$ so that its i^{th} column, $\mathbf{d}_i$, then represents the "structural shock" to the ith element in $\mathbf{y}_t$.

In our paper, we use the ordering with $\mathbf{y}_t = [\Delta csp_t, \Delta comp_t, g_t, \pi_t, pr_t, \Delta reer_t, dr_t, lr_t]'$. The main goal here is to ensure that the identified monetary policy shocks are exogenous. Hence, the safest route in general would be to include all available information to identify the shock (Jordá 2023). Thus, global resource demand (Δcsp_t) and global commodity prices ($\Delta comp_t$) are included in the system to capture responses to fluctuations in the external environment and ordered as the first two variables in line with the existing SVAR literature for commodity-exporting emerging markets (Gan-Ochir et al. 2023). Small open economy New Keynesian models for commodity-exporting economies (i.e., Bergholt et al. 2019) are used as the theoretical foundation for the ordering of the next four variables ($g_t, \pi_t, pr_t, \Delta reer_t$)[11]. For the ordering of the last two variables (dr_t, lr_t), we rely on the interest rate pass-through literature (i.e., Hristov et al. 2014), suggesting that bank deposit and lending rate rates are accordingly adjusted in response to a change in policy rate. The ordering is also in line with the fact that the bank rates reflect fluctuations in foreign and domestic market variables, including global commodity prices, inflation, and exchange rate (Gan-Ochir et al. 2023).

3.2 Non-linear local projections

Using non-linear local projections, we examine the state-dependent and asymmetric effects of monetary policy shocks.

Threshold local projections: State-dependent effects of monetary policy shocks

We use threshold local projections to test whether the impulse responses after a monetary policy shock vary in different economic states. Following the approach popularized by Ahmed and Cassou (2016), we consider the binary state (i.e., $\mathrm{k} = \mathrm{HI}, \mathrm{LO}$) setup as follows:

$$\mathbf{y}_{t+s} = I_{t-1}\left[\boldsymbol{\alpha}^s_{\mathrm{HI}} + \sum_{i=1}^{p} \mathbf{B}^{s+1}_{i,\mathrm{HI}} \mathbf{y}_{t-i}\right] + (1 - I_{t-1})\left[\boldsymbol{\alpha}^s_{\mathrm{LO}} + \sum_{i=1}^{p} \mathbf{B}^{s+1}_{i,\mathrm{LO}} \mathbf{y}_{t-i}\right] + \mathbf{u}^s_{T,t+s} \qquad (4)$$

[10] As shown in Plagborg-Møller and Wolf (2021), identification methods common in the VAR literature can be easily incorporated.

[11] The IS curve explains why domestic output growth comes before inflation. Taylor rule suggests that the policy rate should be placed after output growth and inflation; and uncovered interest rate parity (UIP) in real terms describes why real exchange rate change is placed in the last. As the exchange rate contemporaneously responds to all variables, the VAR system can also capture how exchange rates reflect developments in global commodity markets.

for $s = 0,1,2, \dots, h$. The threshold dummy variable is denoted by I_t. Since the value of I_t depends on the economic state (i.e., the size of the interested variable relative to its trend value), it is effectively a threshold dummy variable that introduces non-linearity into the model.

In this paper, we consider the following economic states: i) the economy is in a high or a low policy rate, ii) the economy is in a high or a low inflation, iii) the economy is in a recession or an expansion, and iv) joint states.

Business cycles. Following Ahmed and Cassou (2016), we build the threshold dummy variable using an estimate of the business cycle dating (i.e., output gap). The dummy variable I_t is set equal to 1 in a recession and equal to 0 in an expansion:

$$I_t = \begin{cases} 1 & \text{when the economy is in a recession} \\ 0 & \text{when the economy is in an expansion} \end{cases} \tag{5}$$

Inflation regimes. Following Ascari and Haber (2022), we define the economy to be in the high inflation (π_t) regime if $\pi_t > \bar{\pi}_t$, where $\bar{\pi}_t$ is a trend inflation, measured by a twenty-five month centered moving average, and in the low inflation regime if $\pi_t \leq \bar{\pi}_t$. The dummy variable I_t is set equal to 1 in a high inflation regime and equal to 0 in a low inflation regime:

$$I_t = \begin{cases} 1 & \text{when the economy is in a high inflation regime } (\pi_t > \bar{\pi}_t) \\ 0 & \text{when the economy is in a low inflation regime } (\pi_t \leq \bar{\pi}_t) \end{cases} \tag{6}$$

Policy rate cycle. In line with Alpanda et al. (2021), we define the economy to be in the tightening phase of policy rate (pr_t) if $rpr_t > \overline{rpr}_t$, where $rpr_t = pr_t - \pi_t$ is real policy rate, $\overline{rpr}_t$ is a trend policy rate (a proxy of the neutral rate of interest), captured by a twenty-five month centered moving average of the real policy rate, and in the loosening phase of policy rate if $rpr_t \leq \overline{rpr}_t$. The dummy variable I_t is set equal to 1 in a tightening phase and equal to 0 in a loosening phase:

$$I_t = \begin{cases} 1 & \text{when the economy is in a tigtening phase of monetary policy } (pr_t > \overline{pr}_t) \\ 0 & \text{when the economy is in a loosening phase of monetary policy } (pr_t \leq \overline{pr}_t) \end{cases} \tag{7}$$

Joint states. As an example, we consider a scenario involving a recession combined with a tightening phase of the policy rate. Analyzing monetary policy effectiveness in this context is particularly insightful. In Mongolia, this combination has been frequent, driven by balance of payments (BOP) difficulties. Such challenges often trigger a sharp depreciation of the domestic currency, prompting increased dollarization as foreign currencies offer relatively higher returns. In response, the central bank raises the policy rate to maintain the domestic currency's base return and stabilize the financial system. The dummy variable I_t is set equal to 1 in a recession combined with the tightening and equal to 0 in other states:

$$I_t = \begin{cases} 1 & \text{when the economy is in a recession with tigtening of policy rate} \\ 0 & \text{when the economy is in other states} \end{cases} \tag{8}$$

By analogy to (3), the IRFs for the two states of the economy are calculated by

$$\widehat{IR}^{\mathrm{HI}}(t, s, \mathbf{d}_i) = \widehat{\mathbf{B}}_{1,\mathrm{HI}}^{s} \mathbf{d}_i \qquad s = 0,1,2, \dots, h \tag{9}$$

and

$$\widehat{IR}^{\mathrm{LO}}(t,s,\mathbf{d}_i) = \widehat{\mathbf{B}}^{s}_{1,\mathrm{LO}}\mathbf{d}_i \quad s = 0,1,2,\dots,h \tag{10}$$

with normalizations $\mathbf{B}^{0}_{1,\mathrm{HI}} = \mathbf{I}$ and $\mathbf{B}^{0}_{1,\mathrm{LO}} = \mathbf{I}$. The confidence bands for the IRFs of the threshold model are simple extensions of the method used in the linear LP.

Asymmetric effects of expansionary and contractionary monetary policy shocks

In our analysis, we investigate the asymmetric effects of expansionary and contractionary monetary policy shocks, adopting an approach utilized in recent studies (Tenreyro and Thwaites 2016 and Kloosterman et al. 2024). According to this approach, the structural monetary policy shock (ε_t) is decomposed as $\varepsilon_t = \varepsilon_t^+ + \varepsilon_t^-$ where $\varepsilon_t^+ = \max(\varepsilon_t, 0)$ and $\varepsilon_t^- = \min(\varepsilon_t, 0)$ represent positive (contractionary) and negative (expansionary) shocks, respectively. Subsequently, the linear (state-independent) model is modified as

$$\mathrm{z}_{t+s} = \alpha^{s}_{\mathrm{AS}} + \beta^{+}_{s}\varepsilon^{+}_{t} + \beta^{-}_{s}\varepsilon^{-}_{t} + \boldsymbol{\lambda}'_{s}\mathbf{x}_{\boldsymbol{t}} + \mathrm{u}^{s}_{\mathrm{AS},t+s} \quad s = 0,1,2,\dots,h \tag{11}$$

In the analysis, $\mathrm{z}_t = [g_t, \pi_t, pr_t, \Delta reer_t, dr_t, lr_t]'$ is a vector of interested variables, $\mathbf{x}_{\boldsymbol{t}}$ is a vector of control variables that includes one lag of y_t and one lag of the shocks (i.e., ε^{+}_{t-1} and ε^{-}_{t-1}), $\boldsymbol{\lambda}'_s$ is the coefficient vector for $\mathbf{x}_{\boldsymbol{t}}$, and $\mathrm{u}^{s}_{\mathrm{AS},t+s}$ is the error term. In this case, the IRFs for expansionary and contractionary monetary policy shocks are $\{\beta^{+}_{s}, \beta^{-}_{s}\}$. The monetary policy shock (ε_t) is identified from the VAR with the Cholesky decomposition discussed in Section 3.1.

4. Data and empirical analysis

4.1 Data

The dataset (y_t) used in the model estimations includes eight variables for the period 2008M12-2024M9, including two foreign and six domestic variables. In the case of Mongolia, monthly data on bank deposit rate is only available from December 2008.

YoY changes in the logarithm of China steel production (Δcsp_t) serves as a proxy for fluctuations in China's resource demand, as over 90 percent of the total exports of Mongolia, primarily minerals, are directed to China. The indirect measure of China's resource demand is in line with Dungey et al. (2020). We use YoY changes in the logarithm of the global coal prices as a proxy for the growth rate in global commodity ($\Delta comp_t$), as coal exports account for 60 percent of total exports.

Domestic variables include YoY (annual) growth in real GDP (g_t) calculated as the twelve-month difference in the logarithm of seasonally adjusted real GDP[12], YoY (annual) inflation measured as the twelve-month difference in the logarithm of CPI (2015=100) (π_t), policy rate per annum (pr_t),

[12] In the case of Mongolia, real industrial production (IP) data is only available from January 2009, and when calculating YoY growth in real IP, the sample is further shortened to start from January 2010. Therefore, we opt to use calculated monthly real GDP for our main analysis.

the twelve-month difference in the logarithm of REER ($\Delta reer_t$)[13], bank deposit rate per annum (dr_t,) and bank lending rate per annum (lr_t)[14].

China steel production (in tons, csp_t) is observed from Bloomberg and Trading Economics databases. The global coal price (in USD, $comp_t$) is retrieved from the IMF Primary Commodity Price System. Monthly data for policy rate, bank deposit rate, bank lending rate, and REER are directly collected from the Bank of Mongolia's statistical database, and monthly CPI and quarterly real GDP data are taken from the National Statistics Office of Mongolia. To convert quarterly real GDP into monthly real GDP, we first applied the equal allocation method with linear match-sum (dividing the quarterly value by three and assigning it equally to each month). Subsequently, we adjusted the monthly real GDP for seasonality using the X-13ARIMA-SEATS approach.

Linear and nonlinear LP's IRFs are estimated by the method of joint local projection proposed by Jordá (2009).

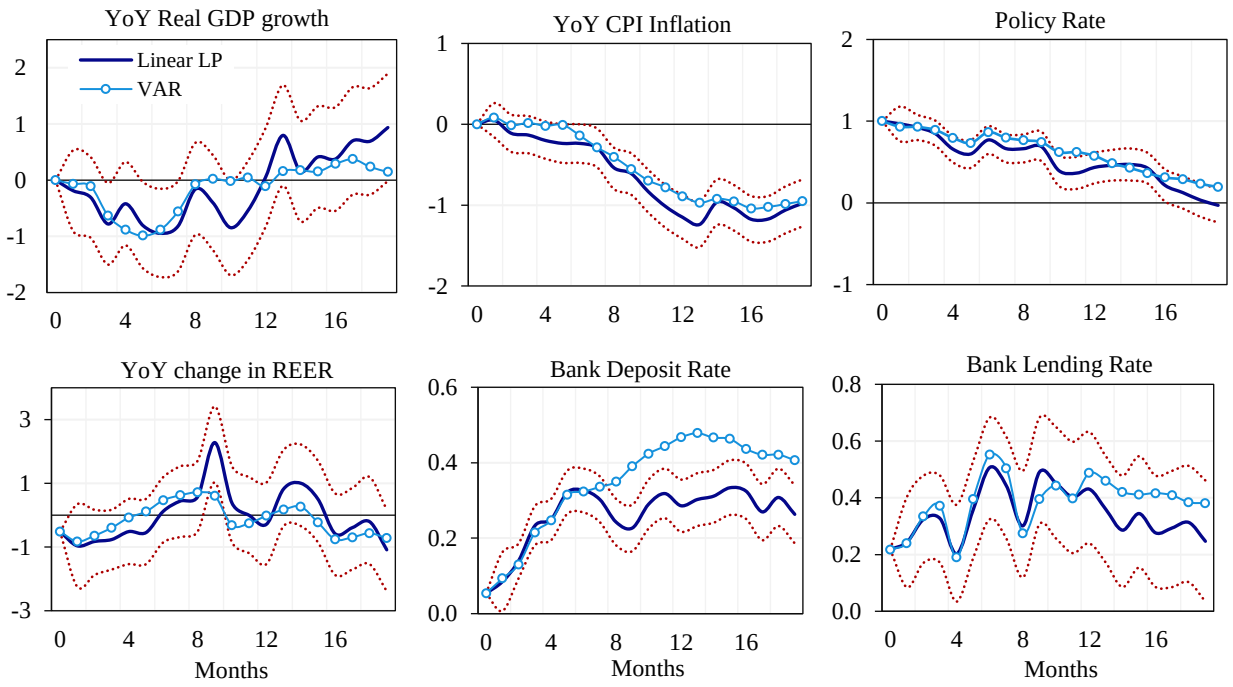
4.2 Linear effects of monetary policy shocks

We begin our analysis by calculating the Impulse Response Functions (IRFs) for the linear LP model presented in equation (2), which serves as the baseline model. Our primary focus is on the results of monetary policy shocks, as our goal is to investigate the monetary transmission mechanism. Based on the Lag Exclusion Wald Tests, we determine the optimal lag length for the system[15]. Given that monthly data is used in the empirical analysis, we select $p = 9$, ensuring no serial correlation and compliance with the stability condition of the VAR system[16]. Figure 3 presents the IRFs to a monetary policy shock, corresponding to a 1 percentage-point increase in the policy rate, for both the linear LP model and the VAR(9) model used the same Cholesky ordering.

There are similarities between the impulse response patterns generated by the linear LP method and the VAR. The responses from the VAR generally align with those from the LP model and remain within the confidence bands, except in the case of the bank deposit rate. However, three key differences emerge. First, the VAR impulse responses are smoother, which reflects differences in the construction methodologies of the two approaches. Second, the declining response of the GDP growth is long-lasting for the LP method, and the response of YoY percentage changes in the REER is statistically significant at the ninth horizon in the case of the LP method. Third, the monetary policy pass-through to the bank deposit rate is estimated to be higher over longer horizons (after eight months) when analyzed using the VAR method.

[13] Increases and decreases in REER indicate appreciation and depreciation of the REER, respectively.
[14] When replacing the REER by the nominal effective exchange rate (NEER) in the estimations, our results remain robust.
[15] In the testing, we include the maximum of 12 lags, and p-values of joint statistics suggest potential lags of 1, 7, 9, and 12 since the null hypothesis of the chi-squared test for lag exclusion is rejected at the 5 percent significance level.
[16] A model with 12 lags is also estimated as part of the robustness check, with the results discussed in Section 4.6.

Figure 3. Impulse responses to a monetary policy shock, in percentage points



Notes: The solid line with "o" marker depicts the point estimates for the VAR model. The solid line depicts the point estimates for the linear LP model, the dashed lines illustrate the 90 percent confidence bands for the linear LP model. The impulse responses are depicted over a twenty-month horizon.

Based on the linear LP results, the monetary policy shock reduces annual GDP growth for the first 12 months and the decline in annual inflation begins in the third month. The peak significant impact on annual GDP growth occurs in the seventh and tenth months, with a decline of 0.85 percentage points. The maximum response of the annual inflation is recorded in the thirteenth month, showing a decline of 1.2 percentage points. The transmission lag and the maximum effect of monetary policy on inflation are in line with the existing empirical facts (i.e., Havranek and Rusnak 2013 for post-transition economies, and Gan-Ochir and Davaasukh 2023 for Mongolia). We observe a delayed overshooting response of the YoY percentage change of the REER, with a 2.1 percentage point appreciation, statistically significant, occurring in the ninth month. While the estimated interest rate pass-throughs exhibit some persistence, the maximum effects of the policy rate shock on bank interest rates are observed in the seventh horizon. At the horizon, the bank deposit rate increases by 0.34 percentage points, and the bank lending rate rises by 0.5 percentage points. These results align with the short-run interest rate multipliers reported by Gan-Ochir (2016).

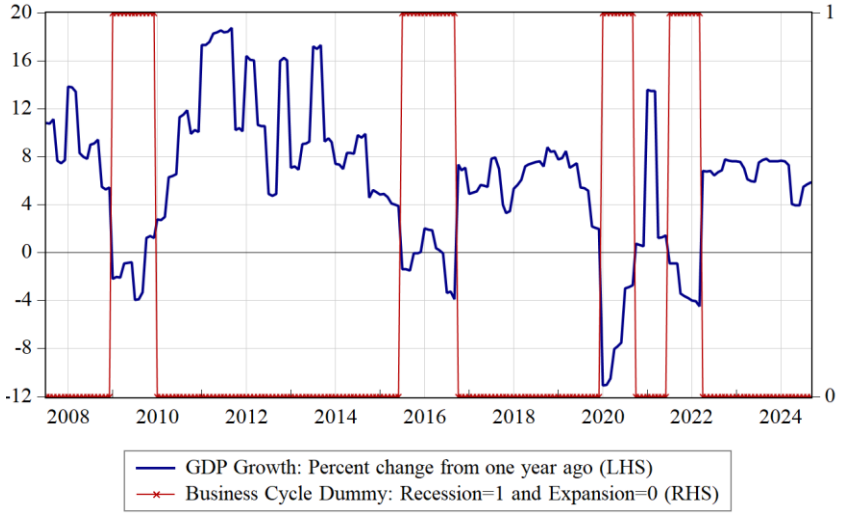
4.3 State-dependent effects of monetary policy shocks

In the first three sub-sections, we report the results for the state-dependent model in (2) for each cycle and show the state-dependent IRFs to a monetary policy shock normalized to decrease (increase) the policy rate by 1 percentage points on impact. In section 4.3.4, we conduct a joint state analysis.

4.3.1 Business cycles

To estimate the threshold model presented in equation (4), it is essential to identify economic states. Business cycle states are identified by defining recessions as peak-to-trough periods, determined using the Bry-Boschan Quarterly (BBQ) algorithm applied to quarterly data. As a robustness check, we also use real GDP gap estimates derived from the Hodrick-Prescott (HP) filter with a smoothing parameter of λ = 1600. The business cycle dating results are consistent among the methods, and recessions are depicted as shaded areas in Figure 1. Based on these estimates and equation (5), we construct the threshold dummy variable (I_t) for business cycles, displayed alongside monthly YoY real GDP growth in Figure 4.

Figure 4. Annual GDP growth and threshold dummy variable for business cycles

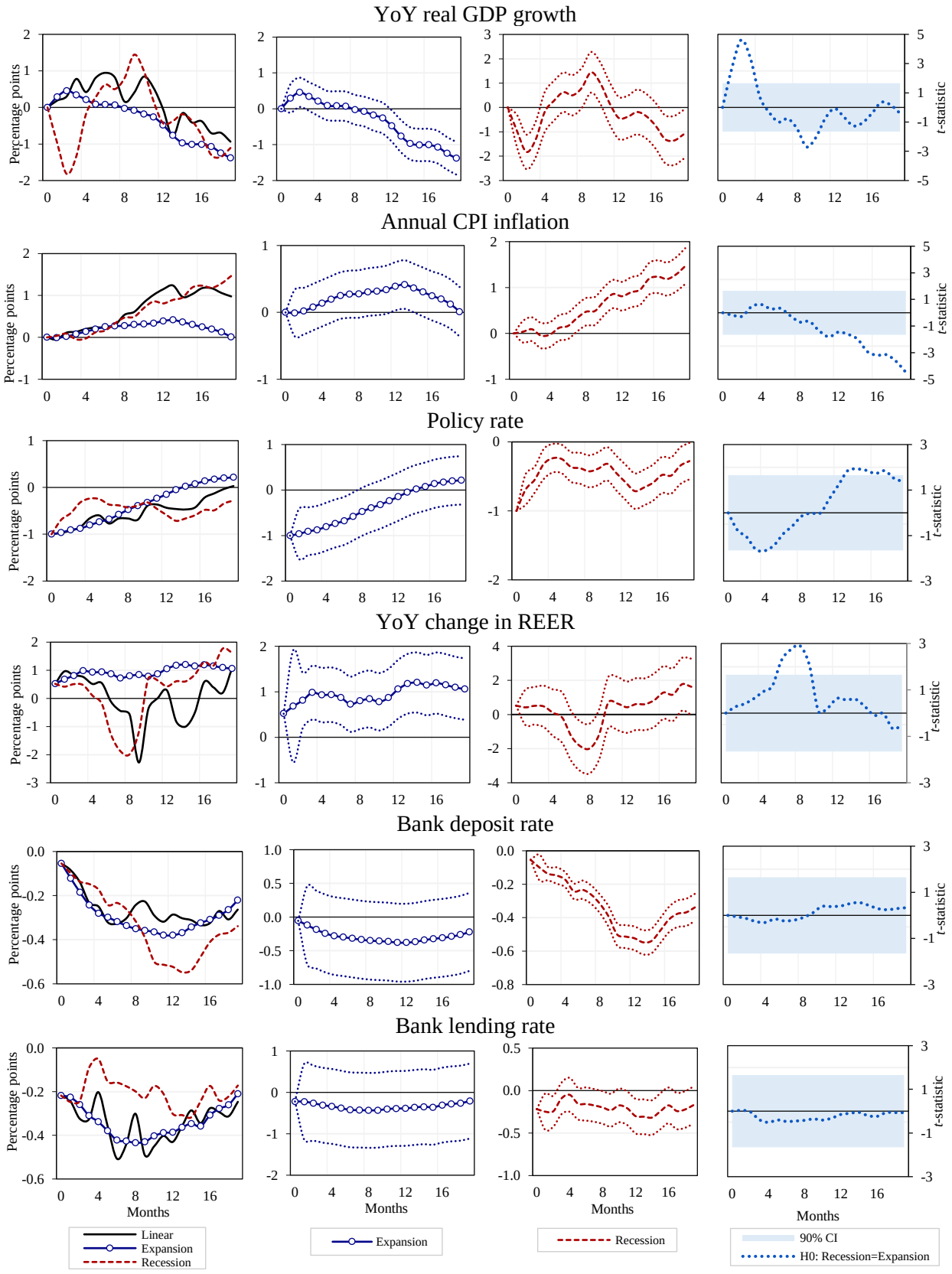


Notes: The solid line plots real GDP growth, measured in percentage change from one year ago on the left axis. The dashed line shows the threshold dummy variable for business cycles on the right axis.

Figure 4 clearly illustrates that the economy experienced recessions during the Global Financial Crisis (GFC), the end of the commodity super cycle in 2015, the COVID-19 pandemic, and China's zero-COVID policy. Year-over-year GDP growth remains negative during most of these recessionary periods.

Figure 5 shows the responses to an expansionary monetary policy shock, corresponding to a 1 percentage-point decrease in the policy rate, across different business cycle phases. The first column displays the point estimates for the linear (solid black), the expansion (solid line with "o" marker) and recession (dashed red line) states. Columns two and three show the impulse responses conditional on the expansion and recession states, respectively, with their 90 percent confidence intervals. The last column depicts the *t*-statistic that tests the null of equality of the expansion and recession state responses (i.e., $\widehat{\mathrm{B}}^{s}_{1,\mathrm{HI}} = \widehat{\mathrm{B}}^{s}_{1,\mathrm{LO}}$), where the light blue area represents the 90 percent *z*-values ([-1.65, 1.65]).

Figure 5. State-dependent impulse responses to a monetary policy shock: Business cycles



The response of annual GDP growth varies significantly across states. During expansions, a positive and statistically significant response of 0.34 percentage points is observed in the third month. In contrast, during recessions, GDP growth initially declines but then rises sharply, exceeding 1.4 percentage points by the ninth month. Additionally, the last column indicates that the different responses of GDP growth in expansionary and recessionary states are statistically significant in short horizons. For annual CPI inflation, the responses are initially similar across states but diverge at longer horizons. During recessions, the responses become statistically significant after seven months and continue to increase over time, similar to the pattern observed in the linear estimate. The *t*-statistic test indicates that the responses in the two states are statistically significantly different at longer horizons. The policy rate response is generally consistent across states. However, during recessions, the response is more persistent at longer horizons. The delayed overshooting response of the percentage change in REER is more pronounced during recessions. The responses of bank deposit and lending rates are statistically significant only during recessions, with the interest rate pass-through estimated at 0.52 for the bank deposit rate and 0.30 for the bank lending rate in the twelfth month. However, the point estimates of interest rate pass-through across the two states are not statistically significantly different.

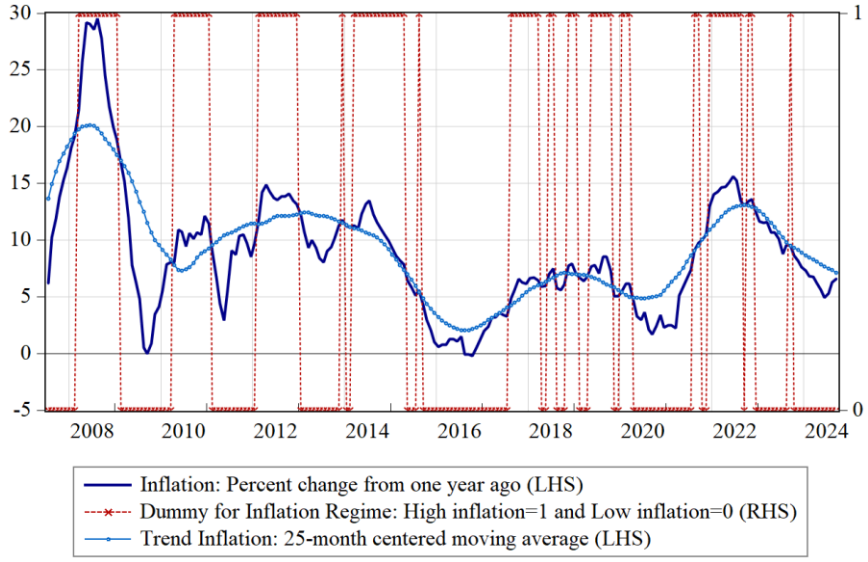
These results point to a stronger effect of monetary policy during recessions. This is consistent with the results found in the existing studies (e.g., Lo and Piger 2005, Santoro et al. 2014). The macroeconomic literature suggests that theories such as convex aggregate supply, the financial accelerator (Peersman and Smets 2005), and the behavioral mechanisms underlying loss-averse preferences (Santoro et al. 2014), can explain why the effects of monetary policy on output are stronger during contractions.

4.3.2 Inflation regimes

Following Ascari and Haber (2022), we measure trend inflation using a 25-month centered moving average (MA). Using equation (6), we then construct the threshold dummy variable (I_t) to represent inflation regimes, which is displayed alongside inflation and its trend measure in Figure 6.

Figure 6 shows that the economy experienced periods of high inflation during several key events: the global Great Recession in 2008, marked by surging global oil and food prices; the economic overheating in 2010-2012, fueled by FDI inflows to the mining sector; the years 2014-2015, characterized by the central bank's quasi-fiscal operations; the economic recovery driven by the IMF's program implemented between 2017 and 2020, and the global surge in inflation in 2022 following the COVID-19 pandemic and the Russia-Ukraine war.

Figure 6. Annual inflation and threshold dummy variable for inflation regimes

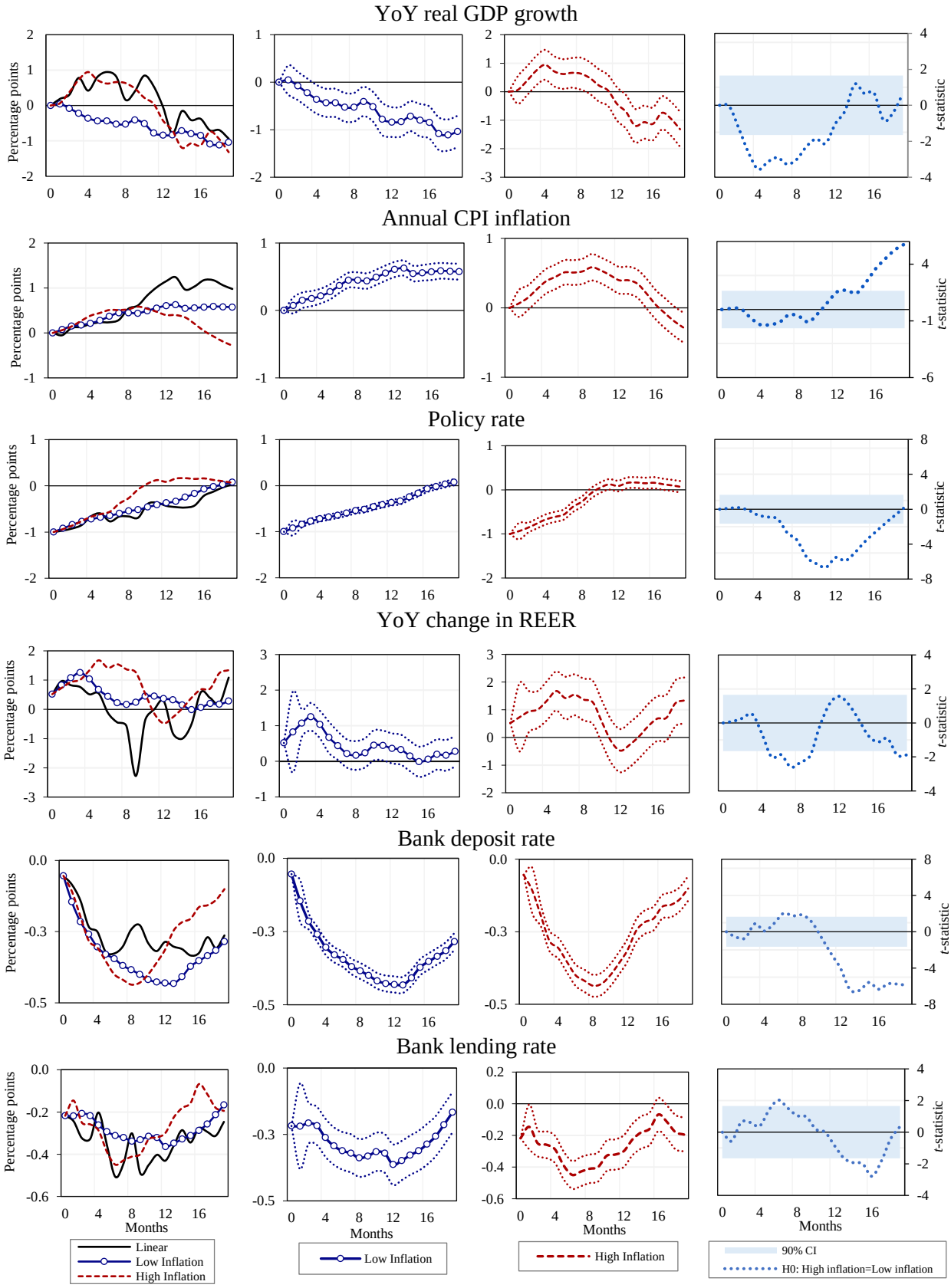


Notes: The solid line plots the annual inflation, measured in percentage change from one year ago, on the left axis. The solid line with "o" marker presents the trend inflation on the left axis. The dashed line with the cross marker shows the threshold dummy variable for inflation regimes, measured on the right axis.

Figure 7 illustrates the responses to an expansionary monetary policy shock across different inflation regimes. The impulse responses for the linear (solid black), the low inflation (solid line with "o" marker), and the high inflation (dashed red line) are reported. Other notations in the figure are identical to those described earlier.

The GDP growth response differs significantly across regimes: during the low inflation regime, the response is negative, while during the high inflation regime, growth rises for twelve months, similar to the linear case. It is in line with the results obtained by Jordà (2020). For CPI inflation, the responses during the high inflation regime are slightly larger than those in the low inflation regime for the first ten months. In the high inflation regime, the peak response of inflation is 0.59 percentage points, occurring in the ninth month. However, at longer horizons, the response is smaller in the high inflation regime. The responses of inflation in the high and low inflation regimes are statistically significantly different at longer horizons. These findings are totally consistent with the results obtained by Ascari and Haber (2021), who interpret them as evidence supporting state-dependent price models as key propagation mechanisms for monetary policy shocks. These models predict a faster and less persistent reaction to monetary disturbances in the high inflation regime. The response of the policy rate during low inflation regimes aligns with the linear case, remaining negative for approximately twenty months. In contrast, the policy rate stays negative for only nine months in the high inflation regime. These differences in policy rate reactions are statistically significant for horizons between eight and eighteen months. As suggested by Ascari and Haber (2021), the larger responses in the low-inflation regime compared to those in the high-inflation regime also indicate stronger endogenous monetary policy feedback in response to the shock. In both regimes, the response of the change in the REER initially increases, mirroring the response of inflation.

Figure 7. State-dependent impulse responses to a monetary policy shock: Inflation regimes



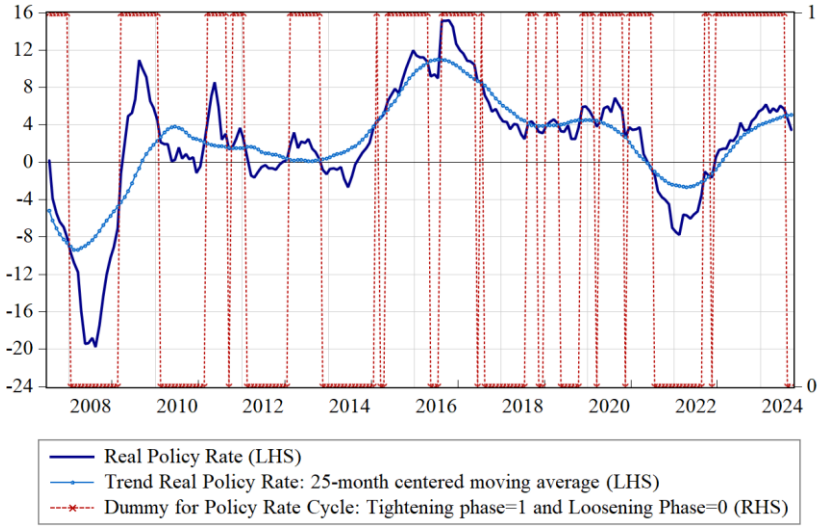
This aligns with the observation that nominal exchange rates are less sensitive to changes in the policy rate and are primarily influenced by the central bank's foreign exchange interventions (FXIs) in the economy. The deposit rate in the two regimes responds differently to a policy rate shock at longer horizons, with these differences becoming statistically significant after the tenth horizon. In the low-inflation regime, the deposit rate initially reacts similarly to the high-inflation regime but increases at a much slower pace thereafter. In the high-inflation regime, the lending rate initially reacts more strongly and then increases much faster than in the low-inflation regime. However, these differences are statistically significant only for four horizons, between the 15th and 18th months. In the high inflation regime, the interest rate pass-through is estimated at 0.5 for the bank deposit rate and 0.43 for the bank lending rate in the seventh month. These results indicate that (i) the effect of monetary policy is stronger in a high inflation regime over the short horizon (up to 10 months), and (ii) the responses of inflation and bank interest rates to a policy rate shock are more persistent in a low inflation regime, indicating stronger monetary policy effects over longer horizons compared to those in a high inflation regime.

4.3.3 Policy rate cycles

In this section, we aim to address the question: does the phase of the policy rate cycle matter when a monetary policy shock occurs?

Drawing on the analyses of Alpanda et al. (2021) and Eichenbaum et al. (2022), we define policy rate cycles using a measure of the interest rate gap, where the trend interest rate (i.e., the neutral rate of interest) is estimated with a 25-month centered moving average (MA). Using equation (7), we construct the threshold dummy variable (I_t) to represent policy rate cycles. This variable is displayed alongside the real policy rate and its trend measure in Figure 8.

Figure 8. Real policy rate and threshold dummy variable for policy rate cycles



Notes: The solid line plots the annual real policy rate, measured in percentage, on the left axis. The solid line with "o" marker presents the trend real policy rate on the left axis. The dashed line with the cross marker shows the threshold dummy variable for policy rate cycles, measured on the right axis.

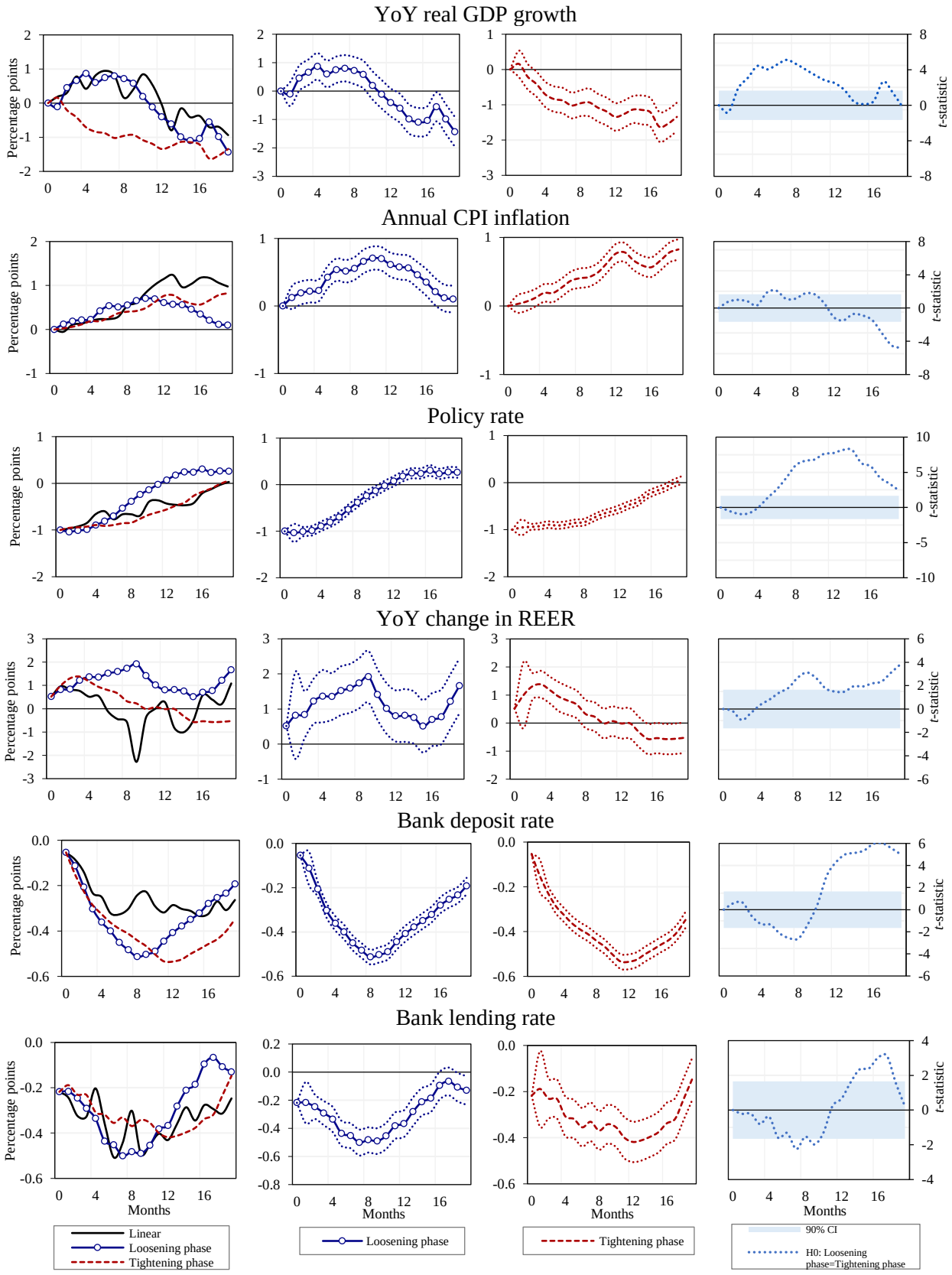
The real interest rate gap in Figure 8 indicates that monetary policy, as reflected in the policy rate, was predominantly tightened during periods of rapid currency depreciation (e.g., 2009, 2015-2016, and 2022) and episodes of high inflation (2021-2023).

Figure 9 displays the responses to an expansionary monetary policy shock, corresponding to a 1 percentage-point decrease in the policy rate, across different (tightening and loosening) phases of the policy rate cycle. The notations in the figure are similar to those described earlier.

The response of the GDP growth differs significantly across phases: during the tightening phase, the response is negative, while during the loosening phase, the response of growth is initially positive, similar to the linear case, peaking at 0.87 percentage points in the fifth month. CPI inflation responses during the loosening phase are initially larger than those in the tightening phase. In the loosening phase, the peak response of inflation is 0.71 percentage points, occurring in the tenth month. However, at longer horizons, the response is smaller in the loosening phase. For instance, in the tightening phase, the response of inflation is 0.78 percentage points in the thirteenth month. The different responses of inflation in the two phases are statistically significant at longer horizons. The response of the policy rate in the tightening phase aligns with the linear case, remaining negative for about twenty months. In contrast, the policy rate response stays negative for only eleven months in the loosening phase. These differences in policy rate reactions are statistically significant for horizons beyond the sixth month. The larger responses in the tightening phase compared to those in the loosening phase also indicate stronger endogenous feedback of monetary policy. In both phases, the response of the change in the REER initially increases. However, in the tightening phase, REER depreciates at the longer horizons. The *t*-statistics test suggests that the stance of monetary policy matters for the responses of change in REER. Bank interest rates in the two phases respond differently to a policy rate shock at longer horizons, with the differences being statistically significant. In the loosening phase, bank interest rates initially react more strongly and then increase much faster than in the tightening phase. In the loosening phase, the interest rate pass-through is estimated at 0.5 for both bank deposit and lending rates in the seventh month. The results indicate that the effect of monetary policy is stronger in a loosening phase of the policy rate cycle over the short horizon.

Overall, the results across the three cycles indicate increased effectiveness of monetary policy in managing inflation when the economy is in a recession, in a high inflation regime, or during the loosening phase with a negative real policy rate gap.

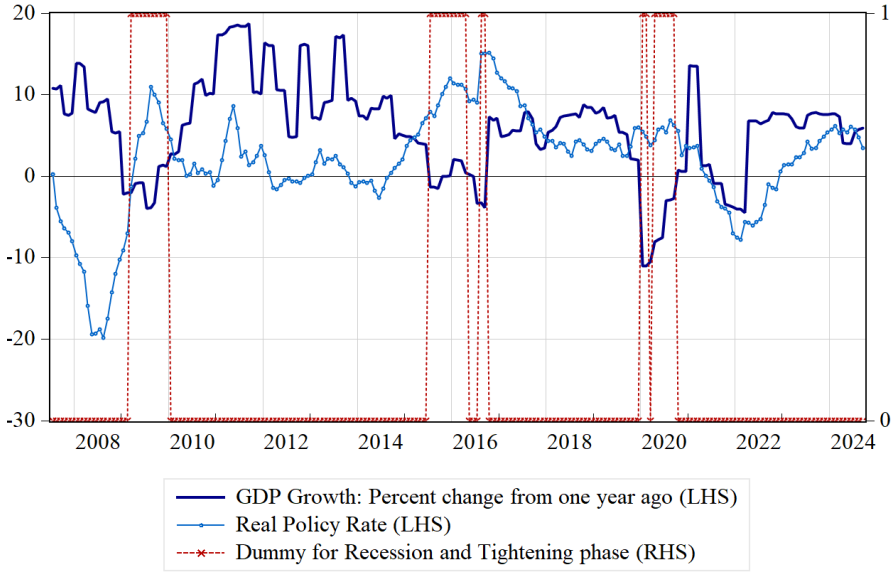
Figure 9. State-dependent impulse responses to a monetary policy shock: Policy rate cycles



4.3.4 Joint state analysis

In this section, we focus on the effectiveness of monetary policy during joint states. Specifically, we investigate whether the impact of policy rate shocks during recessions is influenced by the stance of monetary policy. Using equation (7), we construct the threshold dummy variable (I_t) to represent a recession combined with a tightening phase of the policy rate. This variable is displayed alongside annual GDP growth and the real policy rate in Figure 10.

Figure 10. Real policy rate and threshold dummy variable for policy rate cycles

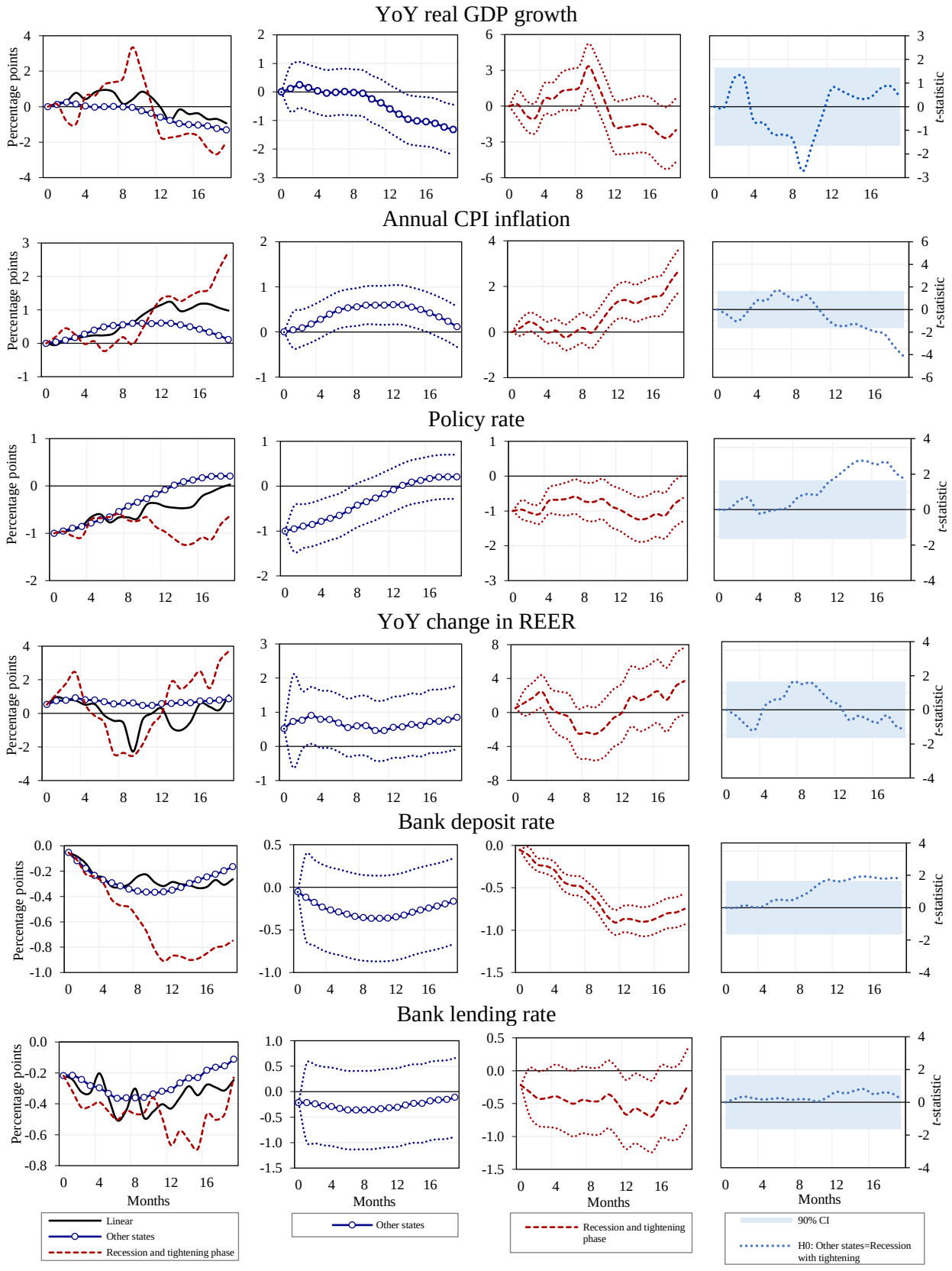


Notes: The solid line plots the annual GDP growth, measured in percentage, on the left axis. The solid line with "o" marker presents plots the annual real policy rate, measured in percentage, on the left axis. The dashed line with the cross marker shows the threshold dummy variable for recessions combined with tightening phases of policy rate, measured on the right axis.

Figure 10 reveals that economic recessions have consistently overlapped with tightening phases of policy rate cycles, except for the most recent recession in 2021. This provides an opportunity to examine the effectiveness of monetary policy during such periods. Figure 11 presents the responses to an expansionary shock across different phases of the joint state (recession with a tightening phase and other states).

The response of GDP growth varies across phases: during other states, the initial responses are flat and statistically insignificant, whereas during a recession with a tightening phase, growth initially increases and peaks in the ninth month. In a recession with a tightening phase, the inflation response is initially statistically insignificant but increases at longer horizons. In contrast, during other states, the inflation response is larger at the short horizon, with a peak of 0.6 percentage points occurring in the tenth month. The inflation responses in the two states are statistically significantly different at longer horizons. The policy rate response in the two states aligns with the linear case over the short horizon. However, during a recession with a tightening phase, the policy rate fluctuates around minus one over the course of twenty months. The differences in policy rate reactions are statistically significant at longer horizons.

Figure 11. State-dependent impulse responses to a monetary policy shock: Joint state



In a recession with tightening phases, a delayed overshooting response of the change in REER is observed. However, the REER responses in the two states are not statistically significant across all horizons. Interest rate pass-through from the policy rate to bank interest rates is larger and statistically significant during a recession with a tightening phase. In this phase, the interest rate pass-through is estimated at 0.87 for the bank deposit rate and 0.67 for the bank lending rate in the twelfth month. Overall, the results indicate that monetary policy is effective in influencing bank interest rates and GDP growth over the short horizon during a recession with a tightening phase.

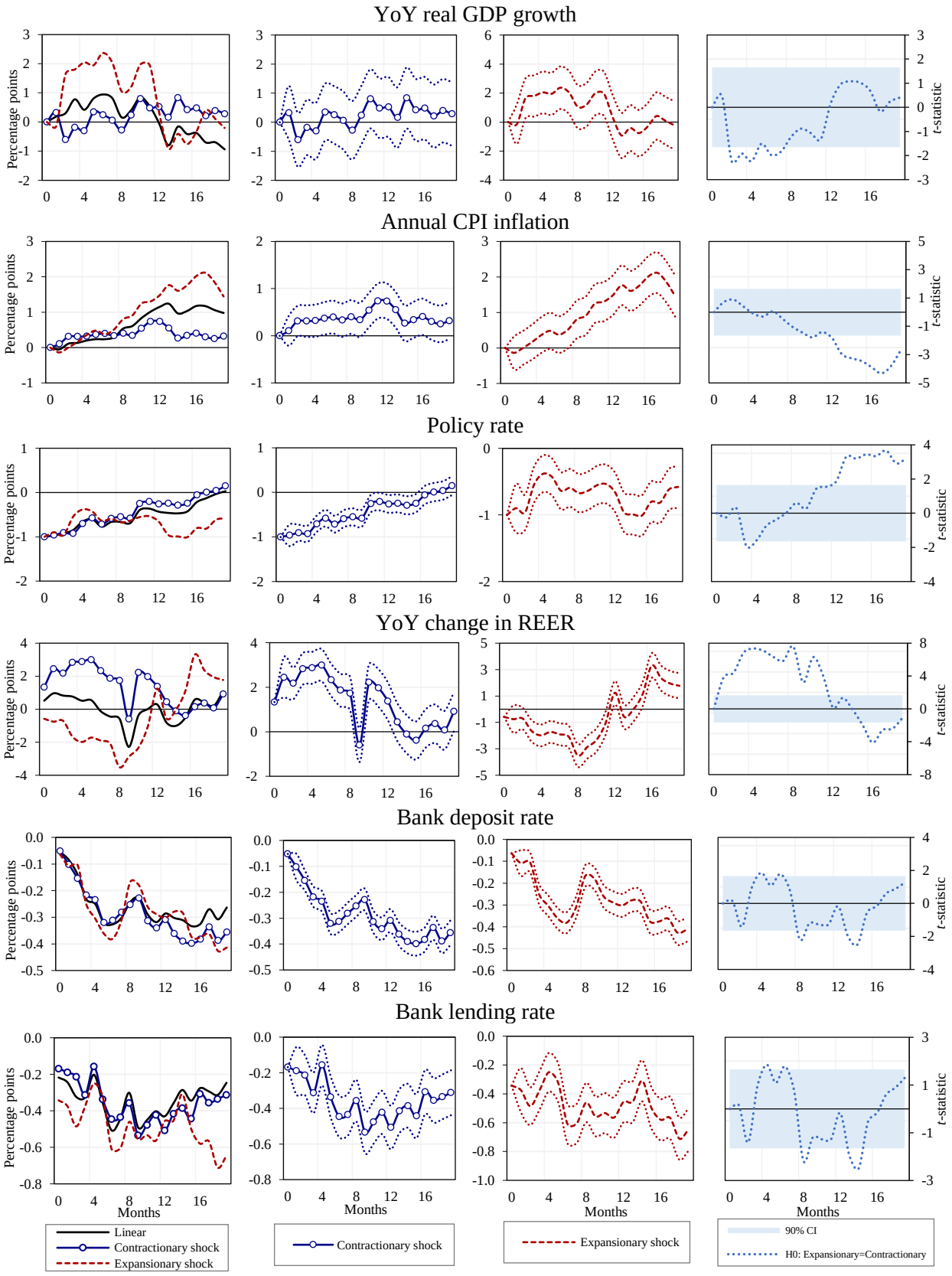
4.4 Effects of expansionary and contractionary monetary policy shocks

In this section, we aim to distinguish the effects of expansionary and contractionary monetary policy shocks. To achieve this, we first extract the structural monetary policy shock from the VAR model (9) using the Cholesky decomposition, as estimated in Section 4.1. Following the methodology outlined in Section 3.2, we calculate the contractionary (ε_t^+) and expansionary (ε_t^-) shocks. Using these series, we estimate impulse response functions (IRFs) based on equation (11).

Figure 12 presents the responses to expansionary and contractionary monetary policy shocks. For comparative purposes, the shocks are normalized to a 1 percentage-point decrease in the policy rate. The notations in the figure are consistent with those described in earlier sections.

Expansionary shocks (i.e., monetary loosening) exhibit a much larger impact on GDP growth compared to contractionary shocks (i.e., monetary tightening). The asymmetric responses of GDP growth are statistically significant at the 10 percent level during the third to eighth horizons, where the response to expansionary shocks remains statistically significant. The effects of expansionary shocks on annual inflation are larger over longer horizons compared to contractionary shocks, with the difference becoming statistically significant beyond the thirteenth horizon. The larger response of inflation to expansionary shocks may be attributed to the persistence of the policy rate response, which remains highly persistent and stays negative for twenty months. The responses of the policy rate to expansionary and contractionary shocks differ statistically significantly over longer horizons. The response of the change in the REER to an expansionary shock exhibits a hump-shaped pattern, indicating depreciation in the REER. Additionally, the responses of the change in REER to expansionary and contractionary shocks are statistically significantly different for most horizons. Expansionary shocks have a larger impact on bank interest rates compared to contractionary shocks, although the responses are not statistically significant for most horizons. In the case of an expansionary shock, the interest rate pass-through is estimated at 0.36 for the deposit rate and 0.6 for the lending rate in the seventh month. Overall, the results suggest that expansionary monetary policy shocks have a stronger impact on key macroeconomic and financial variables compared to contractionary shocks. As argued by Tenreyro and Thwaites (2016), if expansionary shocks were more common during recessions than expansions, the results in Figure 12 might account for the finding that monetary policy tends to be more powerful in recessions. However, no such regime-dependent pattern in the shocks is observed, which is consistent with Figure 10.

Figure 12. Impulse responses to expansionary and contractionary monetary policy shocks



4.5 Robustness checks

In this section, we assess the robustness of our results to different model specifications by (i) changing the sample period, (ii) changing the number of lags, and (iii) modifying the shock identification method in the VAR.

Based on novel and important findings of this paper, impulse responses are chosen in the comparison of alternative specifications. In all cases, the monetary policy shock corresponds to a 1 percentage point decrease in the policy rate. For the linear model, we present the responses of output growth, CPI inflation, and the bank lending rate to the shock (Figure 13A). For the state-dependent model with inflation regimes, the responses of these variables during periods of high inflation are shown (Figure 13B). In the state-dependent model with policy rate cycles, the responses of the variables during the loosening phase are reported (Figure 13C). Finally, for the asymmetric model (i.e., the state-independent model modified for expansionary and contractionary shocks), the responses of the three variables to an expansionary shock are displayed (Figure 13D).

Our key findings remain robust across all alternative specifications. The baseline models are estimated using data from 2008M12 to 2024M9, and their point estimates and confidence intervals are shown as black solid and dashed lines. To examine the impact of the COVID-19 pandemic on the transmission mechanism of monetary policy and test the robustness of our results, we re-estimate the baseline models over a shorter sample period ending in 2020M12[17]. The results, shown as light blue lines with "x" markers in Figure 13, confirm the robustness of our findings. Responses from the shorter sample largely fall within the confidence intervals of the baseline models and generally align with the baseline responses. A slight difference is observed in the state-dependent model with inflation regimes, and the effects of monetary policy shocks on inflation generally appear to strengthen over time.

The baseline models are estimated using nine lags. To test robustness, we re-estimate the models using seven lags, with the results depicted as red lines with "Δ" markers in Figure 13. The responses from these models remain within the confidence intervals of the baseline models.

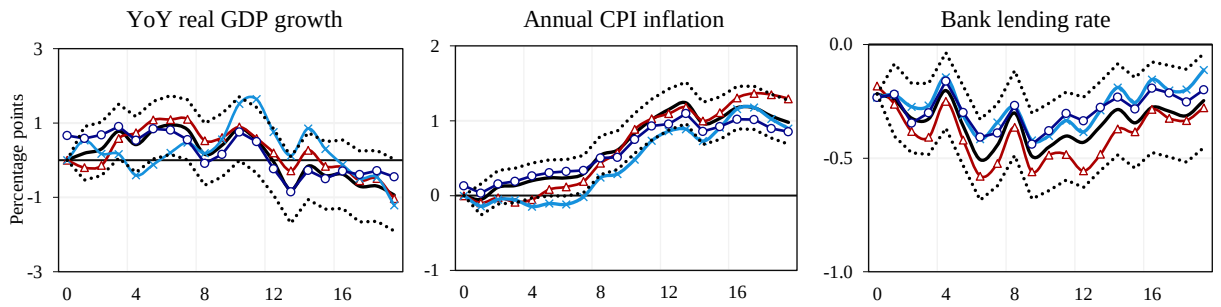
As an alternative shock identification method, we employ the generalized impulse response approach proposed by Pesaran and Shin (1998). Unlike the traditional recursive (i.e., Cholesky) approach, the generalized approach does not require the orthogonalization of shocks and is invariant to the ordering of variables in the VAR. The results from this alternative identification method are shown as blue lines with "o" markers in Figure 13. The responses closely track those of the baseline models.

Overall, the comparative analysis presented in Figure 13 reveals no significant differences across the samples, lag lengths, or shock identification methods.

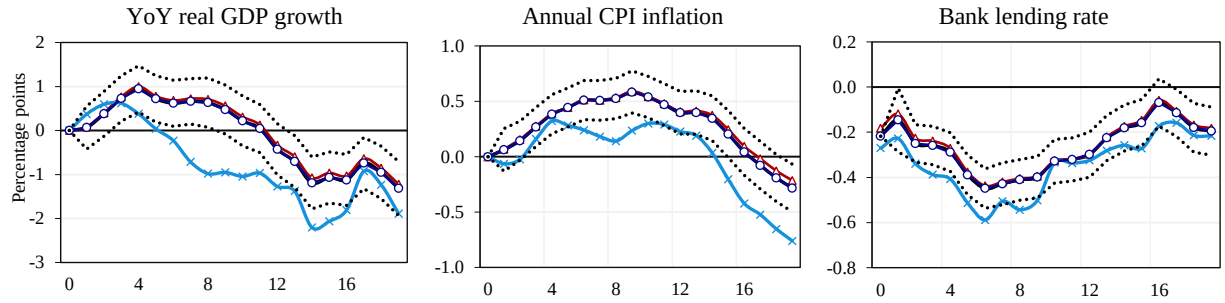
[17] In Mongolia, the first domestic outbreak of the COVID-19 pandemic was recorded on 11 November 2020. Subsequently, strict lockdowns and restrictions were implemented, beginning to impact economic activities.

Figure 13. Impulse responses to monetary policy shocks across different model specifications

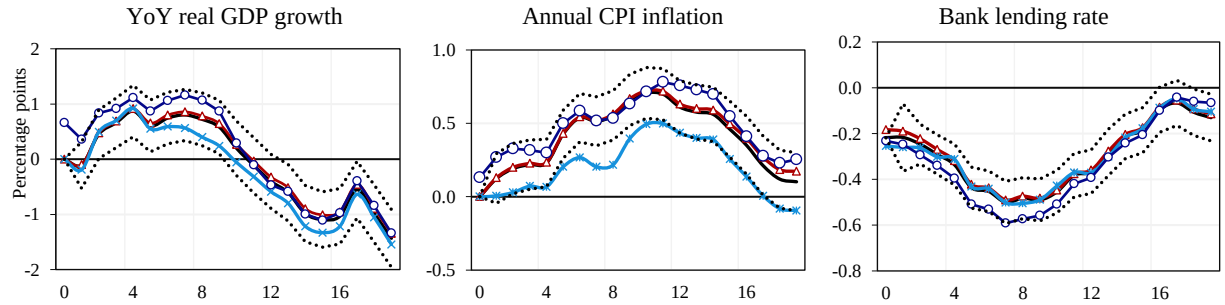
A. Linear model



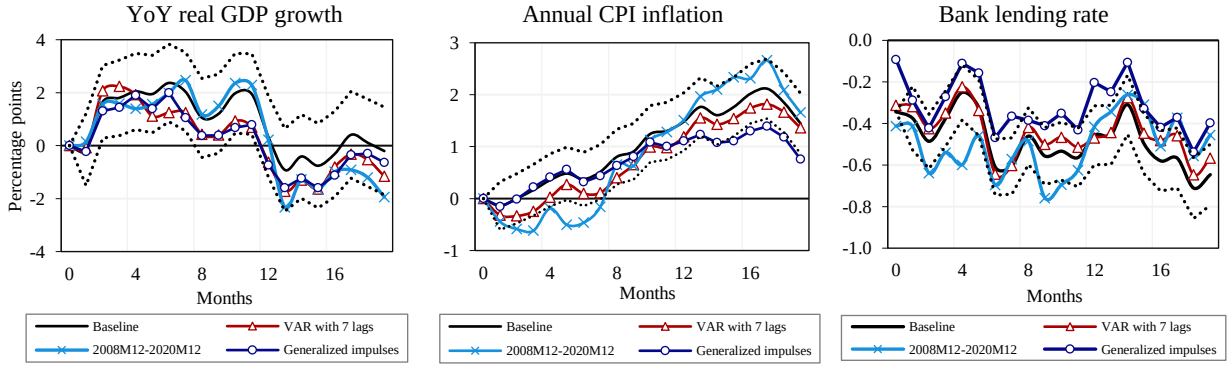
B. State-dependent model with inflation regimes: High inflation regime



C. State-dependent model with policy rate cycles: Loosening phase



D. Asymmetric model: Expansionary shock



5. Conclusion

This paper assessed the effectiveness of monetary policy across the business, inflation, and interest rate cycles, and asymmetric effects of expansionary and contractionary policy shocks in Mongolia, a commodity-dependent developing economy.

This paper provides strong evidence that the effects of monetary policy shocks are both state-dependent and asymmetrical. First, the transmission lag of monetary policy is prolonged and varies with the state of the economy. In the linear model, the peak effect on annual inflation occurs thirteen months after a policy rate shock. Transmission lags are longer in recessions (nine months for GDP growth, twenty months for inflation) compared to expansions (three months for GDP growth, nine months for inflation). Similarly, the transmission lag for the peak effect on inflation is longer under a low inflation regime and during the tightening phase of monetary policy (thirteen months). During recessions, the percentage change in REER exhibits delayed overshooting in response to monetary policy shocks, with the peak effect occurring after eight months. Second, we document stylized facts that transmission channels are sensitive to different phases of the business, inflation, and policy rate cycles. The effectiveness of monetary policy in stimulating GDP is enhanced in economic recessions, monetary loosening phases, high inflation regimes, and the joint state of recession and tightening phase. The ability of monetary policy to manage inflation is diminished during economic expansions. The maximum effects on annual inflation are estimated as 0.42 during expansions, 0.58 in high inflation regimes, and 0.7 during the loosening phase of monetary policy following a shock, corresponding to a 1 percentage point change in the policy rate. Monetary policy pass-through to bank interest rates is stronger during the joint state of recessions and tightening phases but weaker during expansions. For inflation and policy rate cycles, the maximum pass-throughs to bank deposit and lending rates are estimated to be 0.54 and 0.44 on average, respectively. Third, we demonstrate that the transmission mechanism also varies depending on the direction of monetary policy shocks. Transmission lags on inflation are longer for expansionary shocks (sixteen months) than for contractionary shocks (eleven months). Expansionary monetary policy shocks have had stronger effects on key macroeconomic and financial variables than contractionary shocks. Interest rate and exchange rate channels are more pronounced, as expansionary shocks immediately lead to a depreciation of the REER and sharp declines in bank interest rates, boosting GDP growth and inflation. These findings remain robust across various model specifications.

Our analysis highlights the importance of accurately identifying the states of the economy to determine the appropriate timing and magnitude of monetary policy actions. Given that the effectiveness of monetary policy may be constrained in certain circumstances (i.e., economic overheating, the loosening phase, and contractionary shocks), our results suggest that alternative policy tools should be considered during challenging periods. Finally, employing state-dependent models and dynamic policy adjustments can improve responsiveness to both external and domestic shocks, and evolving economic conditions.

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