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

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ABSTRACT

This paper examines the nonlinearity and state-dependence of the monetary policy transmission mechanism in Mongolia—a commodity-dependent developing economy—using local projection methods. Our empirical analysis yields several novel findings. First, transmission lags vary with economic conditions, being longer during recessions compared to expansions. Second, the effectiveness of monetary policy in stimulating GDP is enhanced during recessions, periods of monetary loosening, and high inflation regimes, whereas its capacity to control inflation is reduced during expansions. Monetary policy pass-through to bank interest rates is more pronounced in periods combining recession and monetary tightening. Third, expansionary monetary policy shocks have stronger effects, 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 accounting for economic states when assessing the effectiveness of monetary policy.

1. Introduction

Understanding how monetary policy affects the real economy remains a core concern in macroeconomic research and policy-making. The New Keynesian framework has long provided the theoretical foundation for analyzing the monetary transmission mechanism, emphasizing price rigidities and the role of monetary policy in stabilizing output and inflation (e.g., Mishkin, 1996; Bernanke and Gertler, 1995; Beyer et al., 2024). However, a growing body of empirical evidence (i.e., Ascari and Haber, 2022 for U.S. data, Alpanda et al., 2021 for 18 advanced economies, Berger et al., 2021; Eichenbaum et al., 2022 for U.S. data) has shown that the transmission of monetary policy is state-dependent – its effects vary systematically with the prevailing macroeconomic conditions. Despite these developments, most empirical studies to date have focused on advanced, closed, or financially mature economies, often overlooking the unique structural features of emerging market and developing economies (EMDEs). These economies are typically more exposed to external shocks, have shallower financial systems, and tend to face challenges such as dollarization, commodity dependence, and limited monetary policy credibility, all of which can significantly alter the effectiveness and timing of monetary transmission. In particular, commodity-dependent EMDEs remain underrepresented in this literature, despite their heightened vulnerability to both domestic and external macroeconomic fluctuations.

This paper addresses this gap by studying Mongolia, a small, open, and commodity-exporting EMDE with significant exposure to

☆ The opinions expressed herein are solely those of the authors and do not necessarily reflect the official views of the Bank of Mongolia.

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external shocks and structural constraints. Using the local projection method (Jordá, 2005), this paper empirically investigates how the effects of monetary policy shocks differ across economic states—specifically the business cycle, inflation regimes, and the monetary policy stance—and whether these effects exhibit asymmetries depending on the direction of the shock (tightening vs. loosening). This flexible methodological framework enables a more realistic characterization of how monetary policy operates under varying macroeconomic conditions.

The empirical analysis focuses on Mongolia, a country that encounters significant challenges in managing the effects of external shocks and improving the effectiveness and framework of its monetary policy. Hence, the findings from this case study can be extended to other commodity-dependent EMDEs that face similar structural and macroeconomic challenges.

By focusing on a representative EMDE context, this paper makes two key contributions to the literature. First, it is among the first to systematically examine state-dependent and asymmetric effects of monetary policy across business, inflation, and interest rate cycles in a commodity-dependent developing economy. This enables a comprehensive analysis of how monetary policy influences exchange rate dynamics and interest rate pass-through under different economic states. Second, it provides evidence-based policy insights that can inform more effective and adaptive monetary policy frameworks in commodity-dependent EMDEs confronting similar structural challenges.

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; Álvarez et al., 2019; Bernstein, 2021; Alpanda et al., 2021; Eichenbaum et al., 2022; Kloosterman et al., 2024; Eichenbaum et al., 2025) 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.[1] 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 (i.e., Álvarez et al., 2019), 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, we consider the state dependence of interest rate pass-through along three cycles (i.e., business, inflation, and policy rate cycles). For the existing literature, Puglisi (2015) presents empirical evidence of state-dependent pass-through from the policy rate to lending rates in the US. He argues that a high initial skewness 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 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 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 heterogeneous across countries and types of interest rates and has generally been weaker and slower.

While several studies have examined the transmission of monetary policy in Mongolia, none have employed local projection methods to analyze state-dependent effects. Key findings include asymmetric and delayed interest rate pass-through (Gan-Ochir, 2016), with lending rates responding more to cuts and deposit rates to hikes. Using sign-restricted and Bayesian VARs, recent studies (Gan-Ochir and Davaasukh, 2023; Gan-Ochir, 2023) find that monetary policy shocks affect CPI, GDP, and bank credit with peak effects occurring after 3–4 quarters. Structural models (Gan-Ochir and Undral, 2018; Gan-Ochir and Munkhbayar, 2023, 2024)

[1] Local projection methods are well reviewed and discussed by several papers (i.e., Jordá, 2005; 2023; Plagborg-Møller and Wolf, 2021). 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 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.

estimate a 20 percent pass-through to lending rates and highlight the importance of bank liquidity and credit channels in transmitting policy shocks.

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). According to data from the National Statistics Office (NSO) of Mongolia and the Central Bank of Mongolia (BOM), 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).[2] 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. According to World Bank indicators, Mongolia's export composition closely resembles that of resource-rich EMDEs such as Nigeria, Sudan, Iraq, Turkmenistan, Angola, and Gabon.

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 EMDEs. The import-to-GDP ratio remains high at approximately 65 percent, with imported goods making up nearly 50 percent of the household consumption basket, based on data from the NSO. This level of import dependence is broadly comparable to that observed in other EMDEs such as Georgia, Bulgaria, Moldova, Thailand, Malaysia, and Montenegro. 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. As reported by the BOM, Mongolia's external debt in 2023 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). According to IMF-derived estimations reported by FocusEconomics and CEIC Data, Mongolia's external debt in 2023 substantially exceeds the levels observed in Mauritius (136.1 percent), Senegal (130.1 percent), Laos (126.8 percent), Qatar (123.2 percent), Bhutan (113.0 percent), Venezuela (112.3 percent), and Zambia (105.4 percent).

Mongolia also shares several economic characteristics with MENA and CCA 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).

These structural characteristics common to commodity-dependent EMDEs intensify the presence of evolving real and financial frictions, credit constraints, macroeconomic and financial market volatility, heterogeneous agent behavior, and abrupt shifts in policy environments. From a New Keynesian perspective, such features critically influence the state-dependent nature of monetary transmission mechanisms. In this context, Mongolia presents a compelling empirical case study, offering insights into how these frictions and volatilities shape the effectiveness of monetary policy in a commodity-dependent EMDE setting.

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 (Fig. 1).

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 (Fig. 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).[3]

[2] 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).

[3] 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.

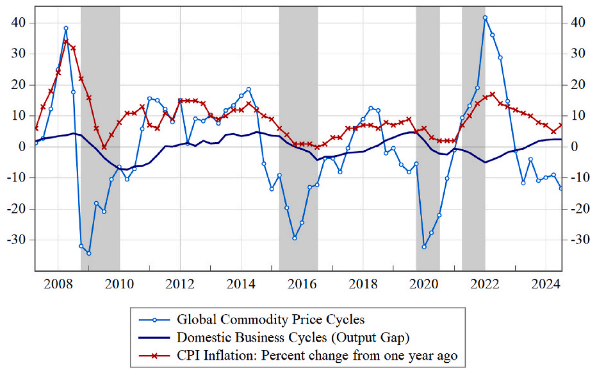

Fig. 1. Global Commodity price cycles, domestic business cycles, and CPI inflation
Notes: Commodity price cycles and domestic business cycles (GDP gap) are calculated in percentage points using the Hodrick-Prescott (HP) filter for quarterly data. The shaded area represents the recession period identified by the Bry-Boschan (BBQ) business cycle dating algorithm.
Source: Authors' calculation using the data retrieved from the NSO of Mongolia and IMF Primary Commodity Prices.

The BOM has been implementing a forward-looking inflation-targeting framework since 2007. As its main policy instrument, the BOM uses the policy rate to signal its monetary policy stance.[4] In July 2007, the BOM introduced the official 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 Fig. 2, when the monetary policy rate was introduced in mid-2007, the bank lending rate was notably high, approximately 22 percent. This was primarily due to a combination of a 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 rate 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

Models grounded in the New Keynesian perspective (e.g., Santoro et al., 2014; Álvarez et al., 2019; Bernstein, 2021; Alpanda et al., 2021; Eichenbaum et al., 2022; Kloosterman et al., 2024; Eichenbaum et al., 2025) suggest that monetary transmission is inherently state-dependent, shaped by evolving nominal rigidities, real frictions, financial market conditions and volatility, the level of inflation, heterogeneous agent behavior, and the surrounding monetary and macroprudential policy environment. Building on these foundational insights, our empirical analysis captures the nonlinear nature of the transmission mechanism by accounting for varying economic environments, such as different phases of business cycles, the prevailing inflation rate, and the stance of monetary policy.

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 (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 small open economy structural VAR models (i.e.,

[4] 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).

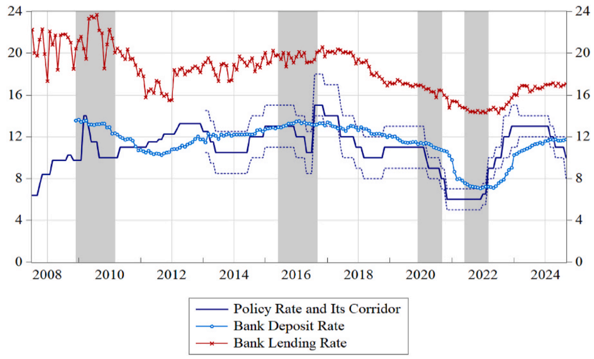

Fig. 2. Interest rates in Mongolia, in percentage points.
Notes: The shaded area represents the economic recession period.
Source: The BOM's Monthly Statistical Bulletin.

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. While the official policy rate is the primary instrument of monetary policy and is included in the system, bank deposit and lending interest rates are also included to reflect changes in broader financial conditions and to assess the effectiveness and transmission of monetary policy through market rates, particularly under different economic states. The inclusion of bank interest rates is further motivated by VAR models used to study interest rate pass-through, such as those by 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:

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

where the operator $E(.|.)$ denotes the best, mean squared error predictor; $\mathbf{y}_t$ is a $n \times 1$ 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-form 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 offers certain advantages over the common vector autoregressive (VAR) models. The LP technique is agnostic about the true DGP[5] 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.[6]

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

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

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 horizons), $\boldsymbol{\alpha}^s$ 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.[7]

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

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

[6] 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).

[7] The maximum lag p can be determined by the information criteria. 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 $\mathbf{X}_t \equiv (\mathbf{y}_{t-1}, \mathbf{y}_{t-2}, \ldots, \mathbf{y}_{t-p})'$.

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

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. Alternatively, 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.

For constructing the experimental shock ($\mathbf{d}_i$), we use methods suggested by Jordá (2005), which follow techniques used in the VAR literature.[8] 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).

Monetary policy decisions often reflect expectations about future economic conditions, raising potential endogeneity concerns when identifying the effects of policy rate shocks. To address this, our empirical strategy combines LPs with shocks identified from a structural VAR using a recursive Cholesky identification scheme. Specifically, we use the following variable ordering: $[\Delta csp_t,\Delta comp_t,g_t,\pi_t,pr_t,\Delta reer_t,dr_t,lr_t]$.[9] The ordering assumes that the policy rate responds contemporaneously to current external and macroeconomic conditions, but not to financial variables such as exchange rates or bank interest rates. This structure helps isolate unanticipated policy rate shocks from endogenous responses to anticipated future developments. The approach is consistent with widely used recursive identification strategies in the monetary policy literature (e.g., Christiano et al., 2005).

In addition, by estimating impulse responses using LPs with control variables, we avoid imposing strong assumptions on the dynamic structure of the data beyond the initial shock identification. The current set of controls is designed to capture external, macroeconomic, and financial market conditions. While incorporating variables that serve as external instruments or proxies for expectations (e.g., high-frequency surprises, narrative shocks, forecast surveys, or forward-looking indicators) would strengthen the identification strategy, the lack of such data in the case of Mongolia limits the feasibility of applying an instrument-based approach such as proxy SVARs or IV-local projection frameworks. To further strengthen the credibility of our identification, we conduct robustness checks using alternative variable orderings and sign-restricted identification schemes. These results are reported and discussed in the robustness checks section.

Given the lack of available data on inflation and commodity price expectations, we partially mitigate the risk of omitted variable bias by including several related observable variables such as global coal prices, China's steel production (as a proxy for external demand), and financial variables including the real exchange rate and bank interest rates, which may indirectly capture market expectations and external influences. We believe the inclusion of these variables helps reduce the potential for bias to a reasonable extent.

3.2. Non-linear local projections

Using non-linear LPs, we examine the state-dependent and asymmetric effects of monetary policy shocks. In this paper, we rely on the threshold local projection (LP) approach for several reasons: i) it allows us to directly test for asymmetries across clearly defined economic states, based on observed gaps in output, inflation, and the policy rate; ii) it offers a clear and intuitive distinction between regimes (e.g., high vs. low inflation or policy rate), which is particularly useful for policy analysis in the context of EMDEs such as Mongolia; and iii) given the relatively short time series available for Mongolia, the threshold approach is more feasible and reliable than smooth transition models, which require larger samples due to greater parameter complexity and nonlinear estimation demands[10].

Following the approach popularized by Ahmed and Cassou (2016), we consider the binary state (i.e., k $=$ HI,LO) setup as follows:

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

[8] 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, $\boldsymbol{\Omega}=\mathbf{PP}'$. 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 i th element in $\mathbf{y}_t$. As shown by Plagborg-Møller and Wolf (2021), identification methods common in the VAR literature can be easily incorporated.

[9] Global resource demand (Δcsp_t) and global commodity prices ($\Delta comp_t$) are ordered as the first two variables in line with the existing SVAR literature for commodity-exporting emerging markets (Gan-Ochir, 2023). Small open economy New Keynesian models for commodity-exporting economies (e.g., Bergholt et al., 2019) are used as the theoretical foundation for the ordering of the next four variables ($g_t,\pi_t,pr_t,\Delta reer_t$). For the ordering of the last two variables (dr_t,lr_t), we rely on the interest rate pass-through literature (e.g., Hristov et al., 2014). 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, 2023).

[10] While smooth transition models offer a valuable alternative for capturing gradual, state-dependent responses, we consider this a promising direction for future research, particularly as longer time series data for Mongolia become available.

for $s = 0, 1, 2, \ldots, h$. 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 a high inflation (π_t) regime if $\pi_t > \overline{\pi}_t$, where $\overline{\pi}_t$ is a trend inflation, measured by a twenty-five month centered moving average, and in the low inflation regime if $\pi_t \le \overline{\pi}_t$. The dummy variable I_t is set equal to 1 in a high inflation regime and equal to 0 in a low inflation regime:

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

Policy rate cycle. In line with Alpanda et al. (2021), we define the economy to be in a tightening phase of policy rate (pr_t) if $r\,pr_t > \overline{rpr}_t$, where $r\,pr_t = pr_t - \pi_t$ is the 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 $r\,pr_t \le \overline{rpr}_t$. The dummy variable I_t is set equal to 1 in a tightening phase and equal to 0 in a loosening phase:

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

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.[11] The dummy variable I_t is set equal to 1 in a recession combined with a tightening and equal to 0 in other states:

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

Given Mongolia's experience with high public debt, quasi-fiscal operations, and directed credit programs, it would be valuable to explore whether fiscal conditions affect the transmission of monetary policy shocks. A promising approach in this regard is Afonso et al. (2025), who analyze the interaction between monetary policy surprises and fiscal regimes in the Euro Area. However, we leave this extension for future research for two reasons: i) estimating fiscal sustainability indicators (e.g., Bohn's coefficients) lies beyond the scope of this paper, and ii) directly adapting their framework to a single EMDE like Mongolia is challenging, as the high persistence of the debt-to-GDP ratio and the relatively short time series lead to an uneven distribution of observations across fiscal regimes. This imbalance reduces the precision of impulse response estimates in the less frequently observed regime, thereby limiting the robustness of regime-specific inferences.

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

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

and

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

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.

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; 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

$$\mathrm{z}_{t+s} = \alpha^s_{\mathrm{AS}} + \beta^+_s \varepsilon^+_t + \beta^-_s \varepsilon^-_t + \lambda'_s \mathbf{x}_t + \mathrm{u}^s_{\mathrm{AS},t+s} \; s = 0, 1, 2, \ldots, h \tag{11}$$

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, $\mathbf{x}_t$ 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}), λ'_s is the coefficient vector for $\mathbf{x}_t$, and $\mathrm{u}^s_{\mathrm{AS},t+s}$ is the error term. In this case, the IRFs

[11] 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.

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. The variables employed in the estimation, along with their respective transformations, are described in detail in the Appendix.

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 prices ($\Delta comp_t$), as coal exports account for 60 percent of total exports of Mongolia.

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), official policy rate per annum (pr_t), 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 official policy rate, bank deposit rate, bank lending rate, and REER are directly collected from the BOM's statistical database, and monthly CPI and quarterly real GDP data are taken from the NSO 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).

All variables used in the empirical analysis, namely, YoY percentage changes ($\Delta csp_t, \Delta comp_t, g_t, \Delta reer_t, \pi_t$), as well as interest rates in levels (pr_t, dr_t, lr_t), are selected based on their theoretical relevance and assessed to be stationary based on both economic reasoning and visual inspection of their time series properties. These transformation choices are consistent with standard practice in empirical macroeconomic research and align with the recommendations of Galí (2015). In particular, the use of YoY inflation rates, interest rates in levels, and log differences of real variables reflects the structure of New Keynesian models, which are typically built on linearized log-deviations from a steady state. In this framework, inflation is measured in growth terms (e.g., annualized or YoY), interest rates are treated as stationary in levels, and real variables are expressed in log differences to capture percentage changes. These transformations help ensure consistency with the theoretical model and mitigate potential issues related to non-stationarity.

We determine the optimal lag length using the Lag Exclusion Wald Test. The test is conducted with a maximum of 12 lags. The p-values of the joint test statistics indicate that lags 1, 7, 9, and 12 are individually significant at the 5 percent significance level. Given the relatively short monthly sample used in the empirical analysis, we select $p = 9$ for all models. This choice ensures the absence of serial correlation in the residuals and satisfies the VAR stability condition. For robustness check, models with 7 lags are also estimated, and the results are presented in Section 4.4.

4.2. Linear and state-dependent 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. Fig. 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.

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. The

[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.

[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.

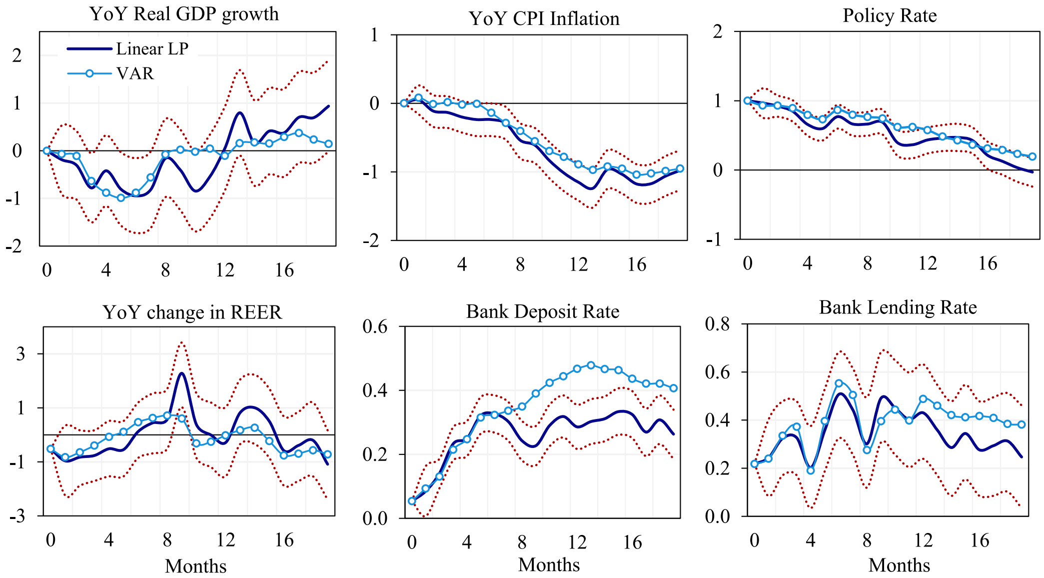

Fig. 3. Impulse responses to a monetary policy shock, in percentage points.
Notes: The solid line with "o" marker depicts the point estimates for the VAR model. The solid line depicts the point estimates for the linear LP model, and the dashed lines illustrate the 90 percent confidence bands for the linear LP model. The impulse responses are depicted over a twenty-month horizon.

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 over-shooting 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).

We now estimate the state-dependent IRFs to a monetary policy shock normalized to decrease (increase) the policy rate by 1 percentage point on impact.

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 $\lambda = 1600$. The business cycle dating results are consistent among the methods, and recessions are depicted as shaded areas in Fig. 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 Fig. 4.

Fig. 4 clearly illustrates that the economy experienced recessions during the Global Financial Crisis (GFC), the end of the

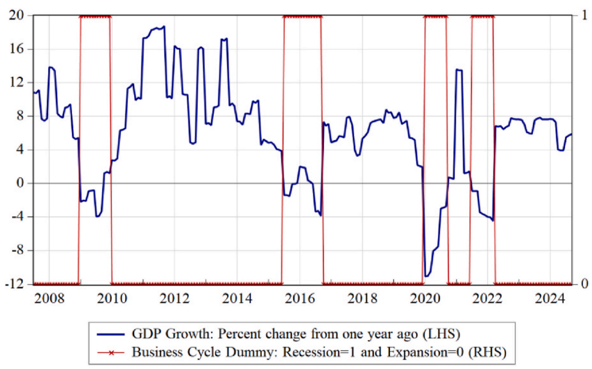

Fig. 4. Annual GDP growth and threshold dummy variable for business cycles.
Notes: The solid line plots real GDP growth, measured in percentage change from one year ago on the left axis. The dashed line shows the threshold dummy variable for business cycles on the right axis.

commodity super cycle in 2015, the COVID-19 pandemic, and China's zero-COVID policy. The GDP growth remains negative during most of these recessionary periods.

Fig. 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), 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., $\widehat{B}^{s}_{1,HI} = \widehat{B}^{s}_{1,LO}$), where the light blue area represents the 90 percent *z*-values ([−1.65, 1.65]).

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.

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.

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 Fig. 6.

Fig. 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.

Fig. 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 et al. (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 (2022), 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 (2022), 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.

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.

Policy rate cycles. Here, we aim to address the question: Does the phase of the policy rate cycle matter when a monetary policy shock occurs?

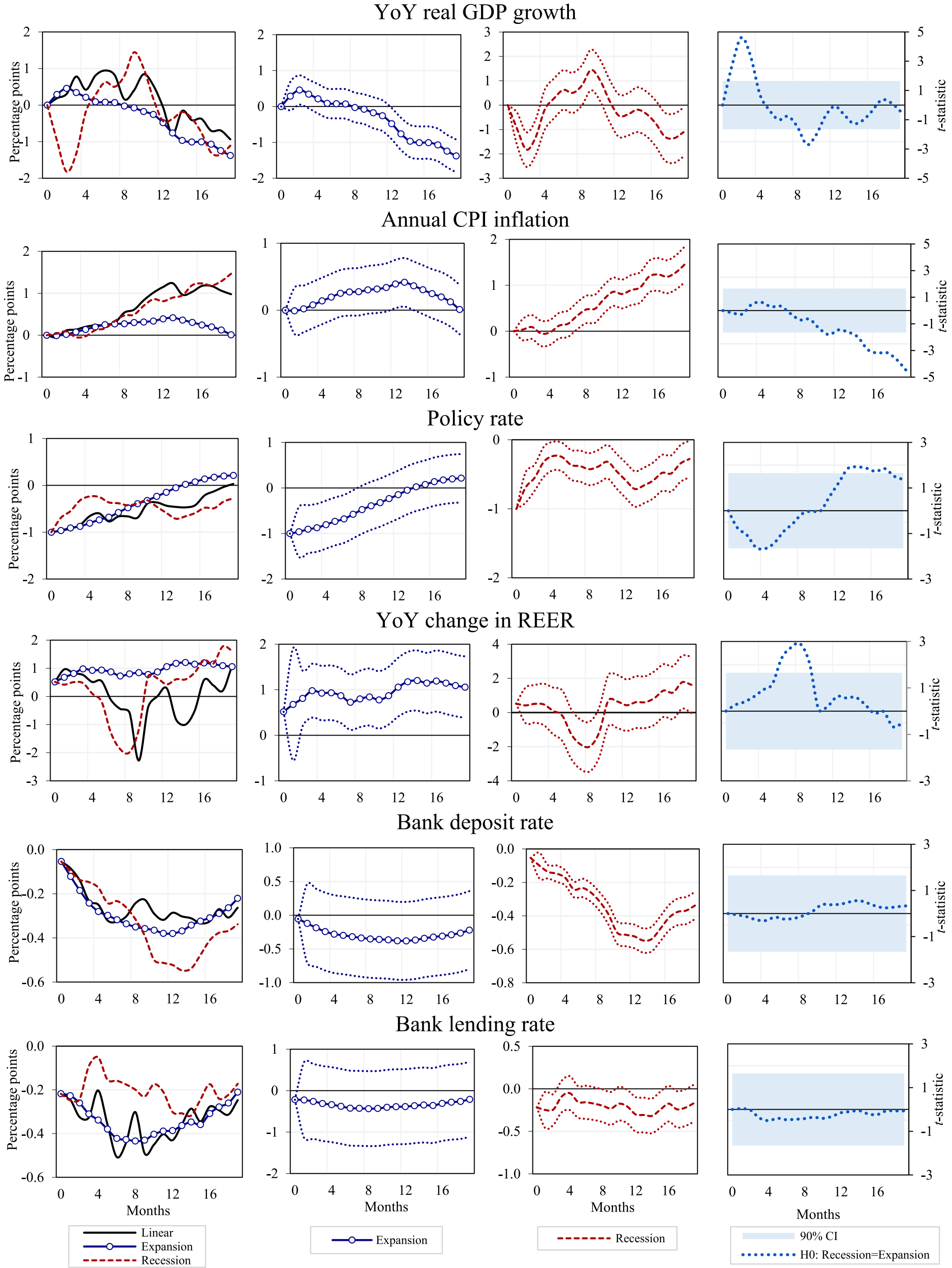

Fig. 5. State-dependent impulse responses to a monetary policy shock: business cycles.

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 Fig. 8.

The real interest rate gap in Fig. 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).

Fig. 9 displays the responses to an expansionary monetary policy shock, corresponding to a 1 percentage-point decrease in the

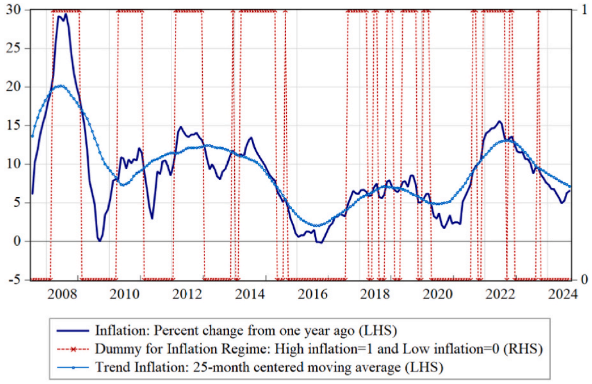

Fig. 6. Annual inflation and threshold dummy variable for inflation regimes. *Notes*: The solid line plots the annual inflation, measured in percentage change from one year ago, on the left axis. The solid line with "o" marker presents the trend inflation on the left axis. The dashed line with the cross marker shows the threshold dummy variable for inflation regimes, measured on the right axis.

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. In the loosening phase, the policy rate response stays negative for only eleven months. 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 changes 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. In line with Eichenbaum et al. (2025), we find that following a period of loosening (tightening) phase in the policy rate cycle, an expansionary monetary policy shock leads to a significant decrease (increase) in the bank's interest rate spread (difference between the lending rate and the deposit rate). Eichenbaum et al. (2025) explain these findings using a nonlinear New Keynesian (NK) general equilibrium model with a banking sector, where some depositors are inattentive to the interest rates earned on deposits. Inattentive depositors may become more attentive after interacting with those who are already attentive. Their NK estimated model shows that these interactions are more likely when interest rates are high.

Joint state analysis. Now, 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 Fig. 10.

Fig. 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. Fig. 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 -1 over the course of twenty months. The differences in policy rate reactions are statistically significant at longer horizons.

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-

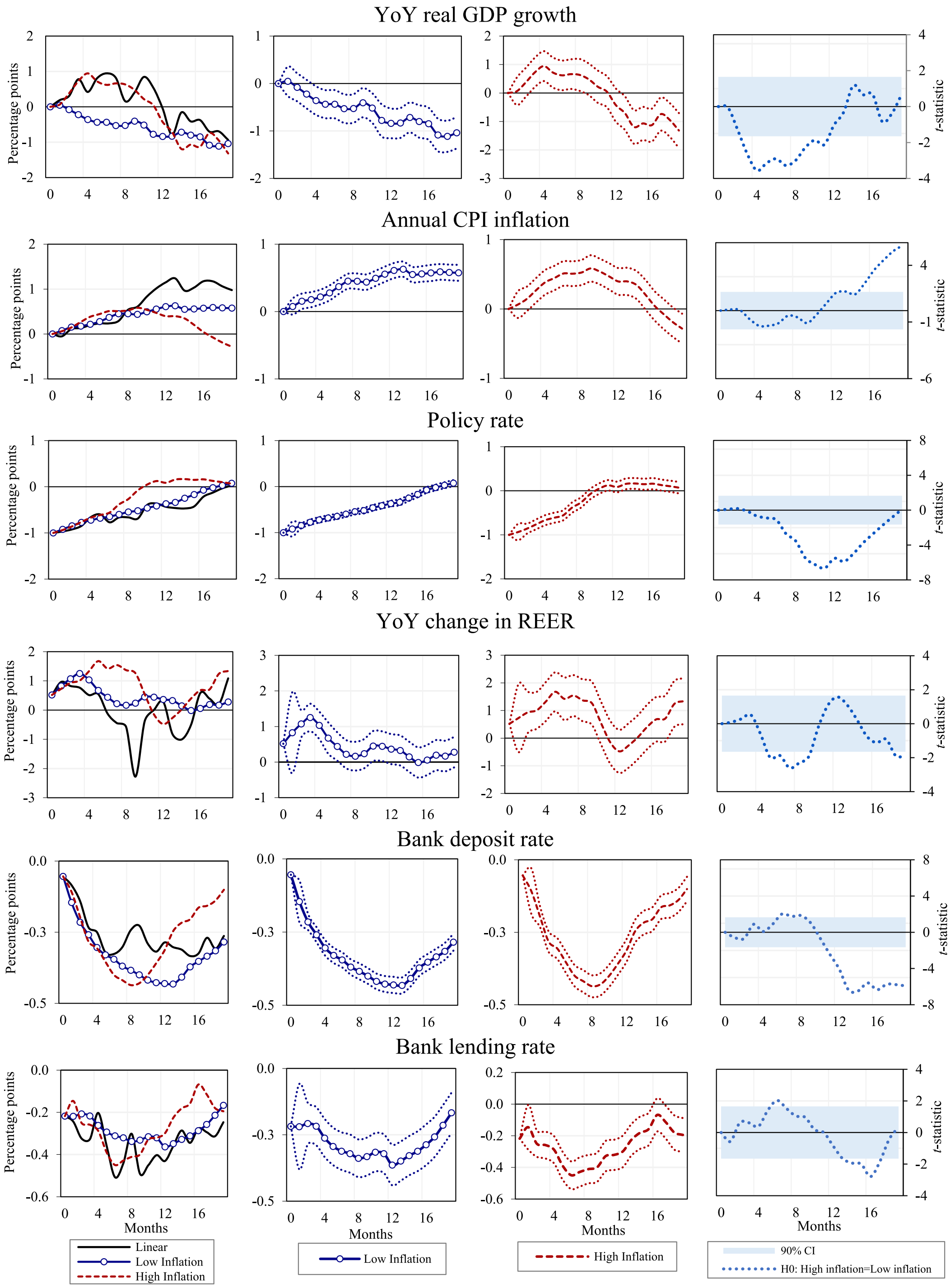

Fig. 7. State-dependent impulse responses to a monetary policy shock: Inflation regimes.

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.

4.3. 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

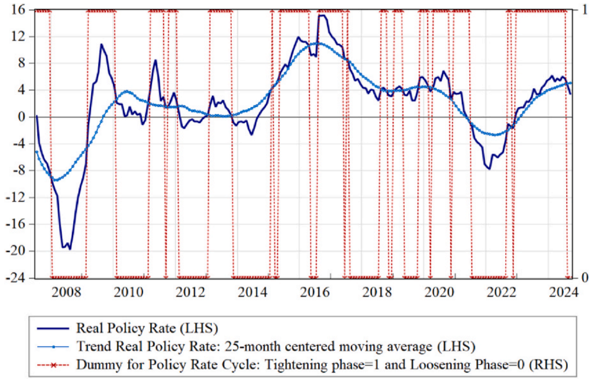

Fig. 8. Real policy rate and threshold dummy variable for policy rate cycles.
Notes: The solid line plots the annual real policy rate, measured in percentage, on the left axis. The solid line with "o" marker presents the trend real policy rate on the left axis. The dashed line with the cross marker shows the threshold dummy variable for policy rate cycles, measured on the right axis.

first extract the structural monetary policy shock from the VAR model (9) using the Cholesky decomposition, as estimated in Section 4.2. 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).

Fig. 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 Fig. 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 Fig. 10.

4.4. Robustness checks

In alignment with the primary objective of this paper, model evaluation focuses on the internal validity of the response functions and the credibility of the identification strategy, rather than out-of-sample predictive performance. As discussed by Ramey (2016), the use of local projections is particularly suited for tracing out responses to identified exogenous shocks, not forecasting, and conventional predictive scoring methods may not align well with causal inference objectives. Moreover, standard cross-validation approaches can be difficult to interpret in settings involving identified shocks (structural VARs or LPs), where the model structure is designed to recover meaningful dynamic responses rather than optimize predictive accuracy (Kilian and Lütkepohl, 2017).

To address robustness, we instead rely on alternative model specifications and verify the stability of the estimated impulse responses across subsamples and identification variants. These approaches are more appropriate for assessing robustness in a causal inference framework. 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 (Fig. 13A). For the state-dependent model with inflation regimes, the responses of these variables during periods of high inflation are shown (Fig. 13B). In the state-dependent model with policy rate cycles, the responses of the variables during the loosening phase are reported (Fig. 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

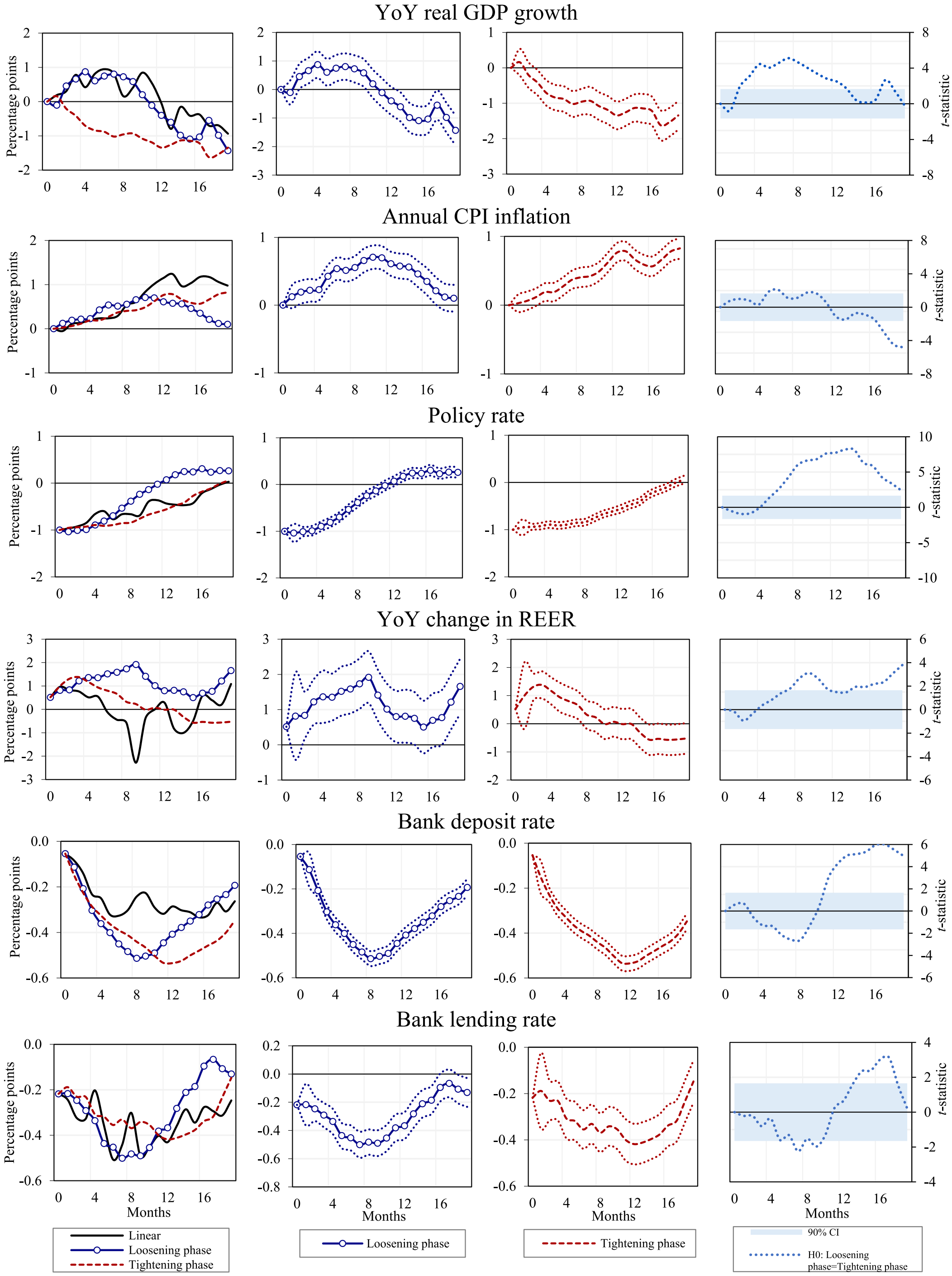

Fig. 9. State-dependent impulse responses to a monetary policy shock: policy rate cycles.

an expansionary shock are displayed (Fig. 13D).

Our key findings remain robust across all alternative specifications. The baseline models are estimated using data from 2008M12–2024M9, and their point estimates and confidence intervals are shown as black solid and dashed lines.

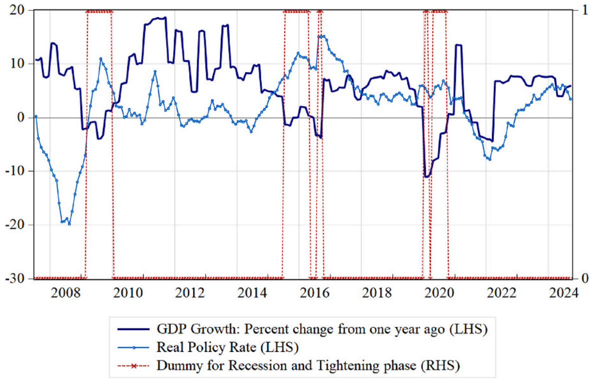

Fig. 10. Real policy rate and threshold dummy variable for policy rate cycles. *Notes*: The solid line plots the annual GDP growth, measured in percentage, on the left axis. The solid line with "o" marker presents the annual real policy rate, measured in percentage, on the left axis. The dashed line with the cross marker shows the threshold dummy variable for recessions combined with tightening phases of policy rate, measured on the right axis.

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.[15] The results, shown as light blue lines with "x" markers in Fig. 13, confirm the robustness of our findings. In the COVID-excluded sample, the estimated responses of GDP growth and inflation are generally weaker across both the linear model and various states of the economy (i.e., the high inflation regime, monetary loosening phase) following expansionary policy shocks, compared to the baseline results. However, the pass-through to bank lending rates appears stronger in the pre-COVID sample, particularly under the high inflation regime and following expansionary monetary shocks. These results suggest that structural shifts and extraordinary events such as the COVID-19 pandemic may have meaningfully altered the effectiveness and transmission dynamics of monetary policy in Mongolia. Nevertheless, most responses estimated from the shorter, COVID-excluded sample fall within the confidence intervals of the baseline models and are generally consistent with the baseline responses. A slight divergence is observed in the state-dependent model with inflation regimes, where the effects of monetary policy shocks on inflation 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 Fig. 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 Fig. 13. The responses closely track those of the baseline models.

To further assess the robustness of our results, we consider an alternative recursive identification strategy for the linear models, in which the policy rate is ordered before inflation (denoted as Alt. SVAR and Alt. LP for the VAR and LP models, respectively). As shown in Fig. 14, the resulting impulse responses are highly consistent with those from the baseline specification, reinforcing the robustness of our findings.

To complement the robustness checks, we estimate a Bayesian Structural VAR (BSVAR) with 9 lags, employing a Minnesota prior, with an overall tightness of 0.01, a relative weight on other variables of 0.99, and a lag decay parameter of 1. The identification of shocks follows the Cholesky ordering used in the baseline model. The resulting impulse responses are presented in Fig. 14. Notably, the BSVAR responses closely align with those from the baseline local projection model, with most estimates falling within the baseline confidence bands. This provides further support for the robustness and stability of our main results.

Overall, the robustness checks reveal no significant differences in the estimated responses across subsamples, lag specifications, or alternative shock identification methods.

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 asymmetric. 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

[15] 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.

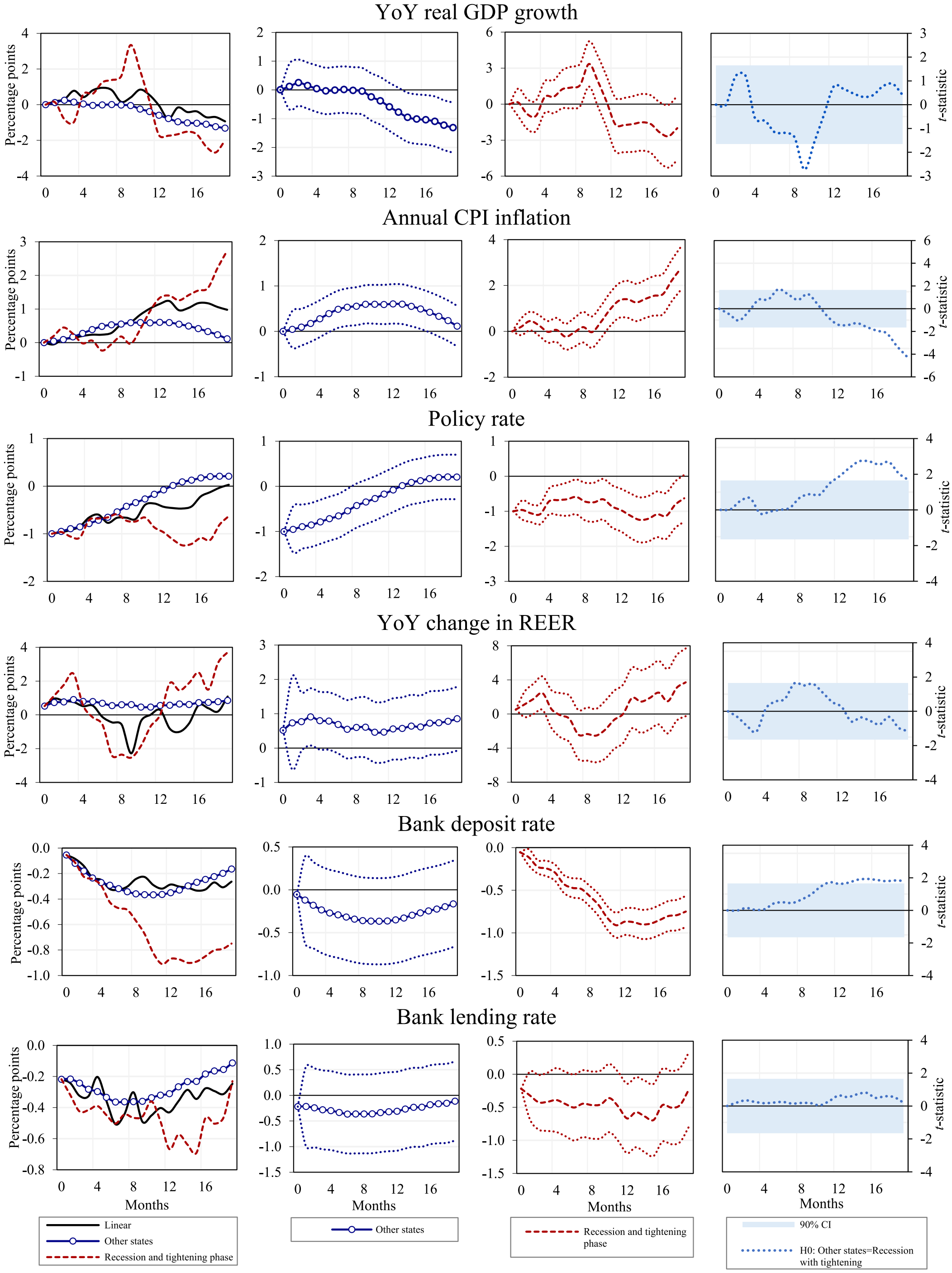

Fig. 11. State-dependent impulse responses to a monetary policy shock: Joint state.

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

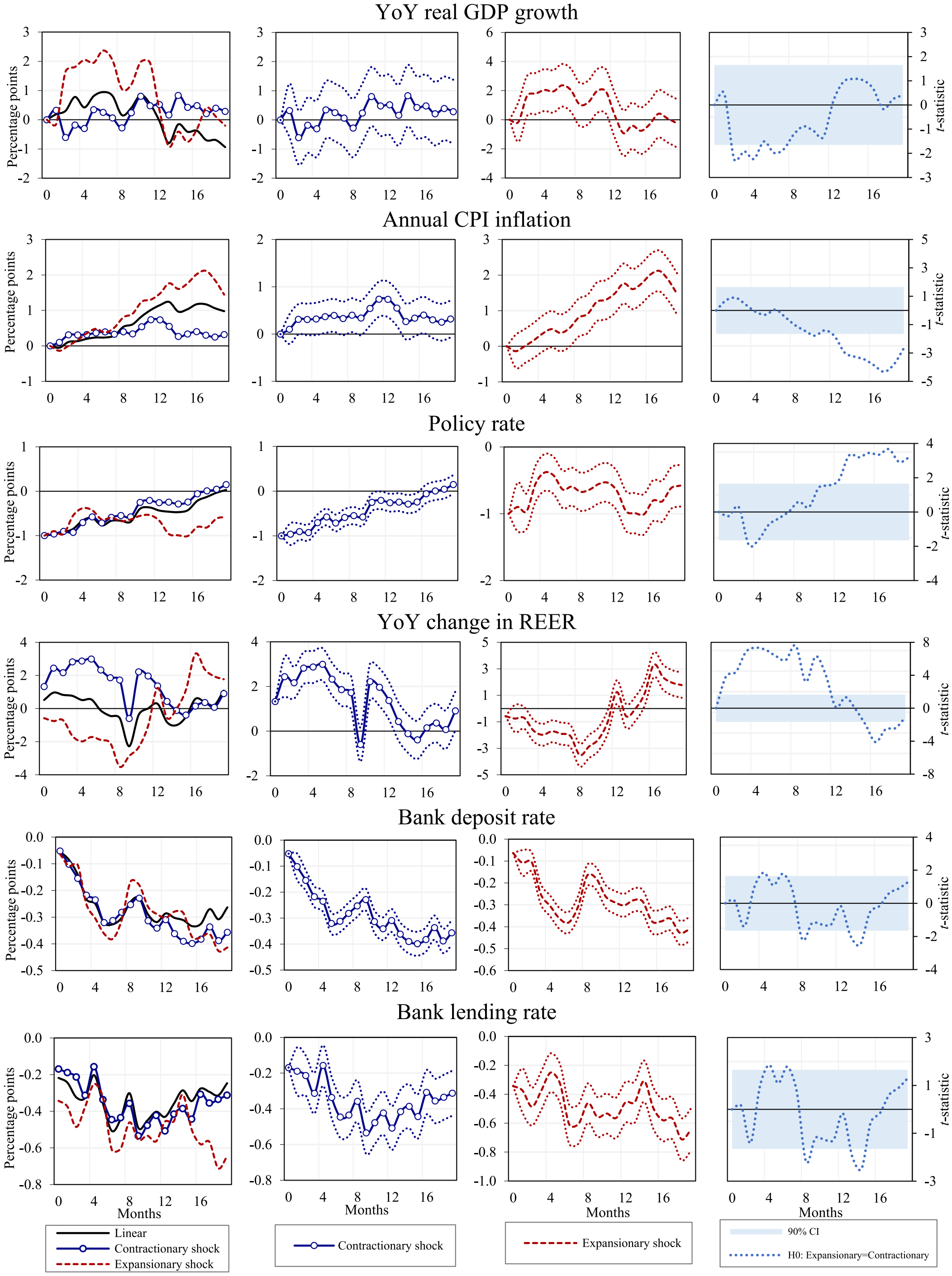

Fig. 12. Impulse responses to expansionary and contractionary monetary policy shocks.

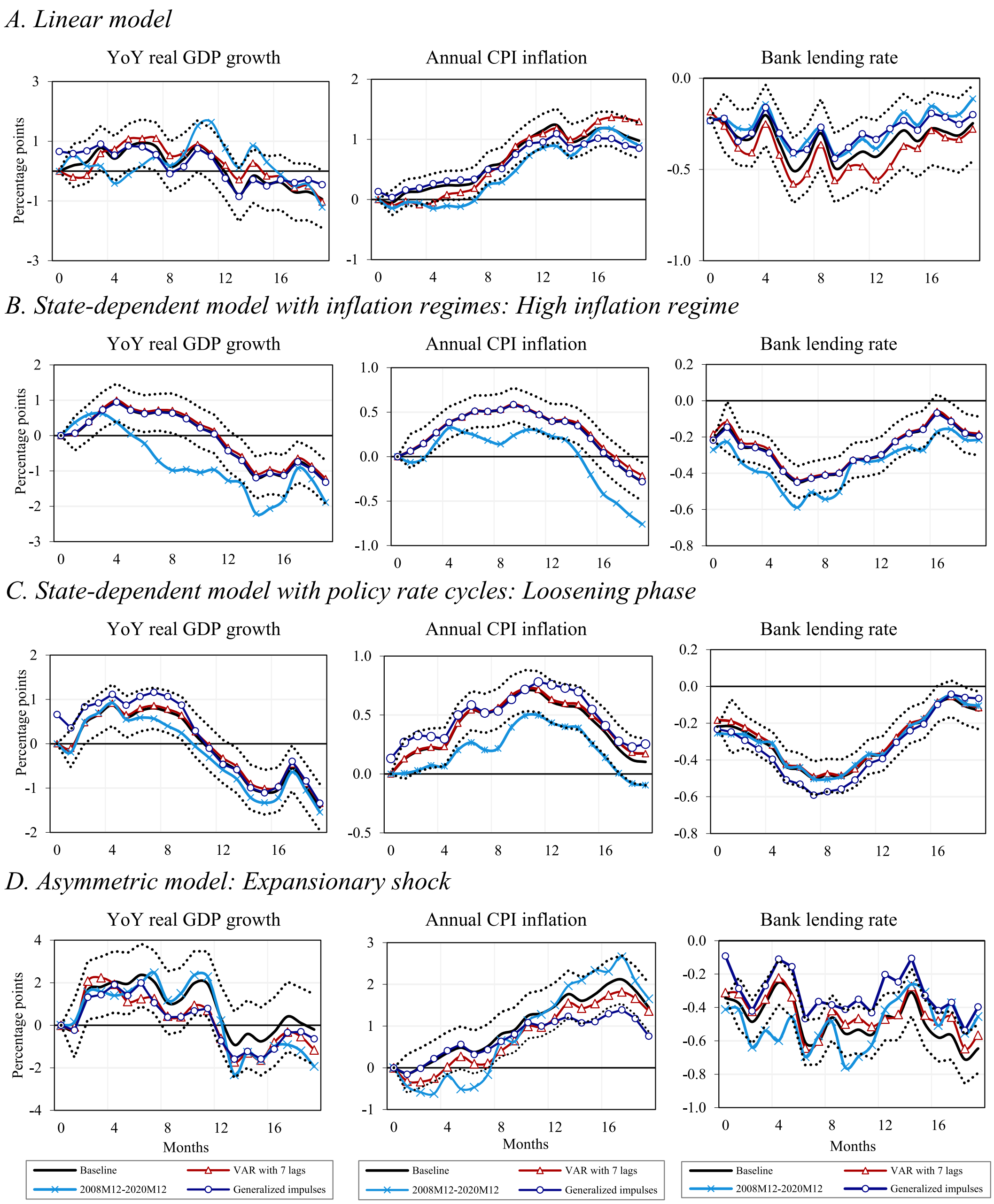

Fig. 13. Impulse responses to monetary policy shocks across different model specifications.

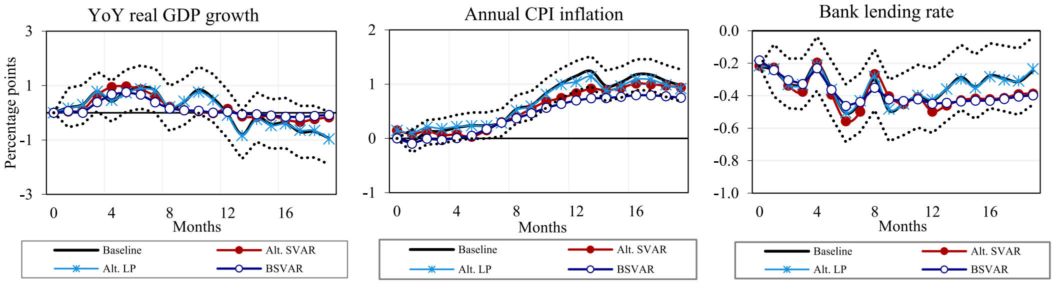

Fig. 14. Impulse responses to monetary policy shocks across different linear models.

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) compared to 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.

Empirical findings from Mongolia offer important preliminary evidence relevant to other EMDEs with comparable structural conditions, namely, high commodity dependence, elevated import reliance, shallow financial markets, significant external indebtedness, and inflation volatility. Conducting a formal comparative empirical study involving other commodity-dependent EMDEs would greatly enhance the external validity and contextual relevance of the results and is recommended as a direction for future research. Another direction for future research would be the development of a stylized New Keynesian model tailored to the structural characteristics of commodity-dependent EMDEs (e.g., commodity price and demand exposure, dollarization, imperfect interest rate pass-through, elevated import reliance, and external indebtedness) to strengthen the narrative around transmission channels. A further extension can explore the use of sign-restricted identification strategies to flexibly capture the dual objectives of inflation targeting and exchange rate stabilization in Mongolia, alongside Bayesian local projections with time-varying parameters to improve robustness and stability of inference.

Our analysis also 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.

CRediT authorship contribution statement

Gan-Ochir Doojav: Writing – review & editing, Writing – original draft, Supervision, Methodology, Formal analysis, Conceptualization. **Arman Juragat:** Writing – original draft, Methodology, Formal analysis, Data curation.

Declaration of competing interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Appendix. Descriptions of data used in the estimation

	Name	Descriptions
Foreign variables	Δcsp_t	The China steel production (csp_t) is observed from Bloomberg and Trading Economics databases (https://tradingeconomics.com/china/steel-production). The year-over-year (YoY) growth rate (Δcsp_t) is computed as the twelve-month logarithmic difference of the series (csp_t).
	$\Delta comp_t$	The global coal price ($comp_t$) is retrieved from the IMF Primary Commodity Price System (www.imf.org/en/Research/commodity-prices). The YoY change in the price is calculated as the twelve-month logarithmic difference of the series ($comp_t$).
Domestic variables	g_t	The real GDP, expressed in 2015 prices, is obtained from the NSO database (https://1212.mn). The real GDP is seasonally adjusted using the X-13 ARIMA-SEATS method, and the quarterly data is converted to monthly frequency using equal allocation via linear match-sum. The YoY growth rate (g_t) is calculated as the twelve-month logarithmic difference of the adjusted monthly series.
	π_t	The consumer price index (CPI), with the base year 2015 = 100, is sourced from the NSO database (https://1212.mn). The annual inflation (π_t) is calculated as the twelve-month logarithmic difference of the CPI series.
	pr_t	The policy rate (pr_t) is directly observed from the Monthly Statistical Bulletin of the BOM (https://stat.mongolbank.mn).
	$\Delta reer_t$	The real effective exchange rate ($reer_t$) is sourced from the BOM (www.mongolbank.mn/en/neer-reer). The YoY change ($\Delta reer_t$) is computed as the twelve-month logarithmic difference of the series ($reer_t$).
	dr_t	The weighted average deposit rate of banks (dr_t) is directly taken from the Monthly Statistical Bulletin of the BOM (https://stat.mongolbank.mn).
	lr_t	The weighted average lending rate of banks (lr_t) is directly taken from the Monthly Statistical Bulletin of the BOM (https://stat.mongolbank.mn).

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