Dynamic Macroeconomic Effects of Monetary Policy in China*

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Abstract

Researchers debate the role of interest rates and credit in implementing monetary policy in China. I compare two alternative proxies to identify China’s monetary policy shocks. The first assumes that the PBOC policy is mainly credit-based and relies on a policy rule for M2 growth. The second uses high-frequency market surprises in interest rate swaps to recover the role of interest rates. Both proxies are compared using an SVAR-IV approach. The peak effects of monetary policy on output and prices are earlier using the credit-based policy proxy. The slower transmission of interest-based policy is robust to alternative specifications, including a novel temporary effect prior. I also provide evidence of how China’s monetary policy transmits to the rest of the world via commodity prices, trade, and financial risk.

Keywords: SVAR-IV, monetary policy identification, China, international monetary policy transmission.

JEL Codes: C32, E52

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1 Introduction

Researchers debate the role of interest rates and credit in implementing monetary policy in China. Chen, Ren and Zha (2018) and Chen, Gao, Higgins, Waggoner and Zha (2020) argue the relevance of the credit channel in implementing the monetary policy stimulus in 2009. They propose an asymmetric policy rule for M2 growth, which is the intermediate monetary policy target. To measure the importance of China in the global financial cycle, Barcelona, Cascaldi-Garcia, Hoek and Van Leemput (2022) use a credit growth measure called credit impulse. Jones and Bowman (2019) argue that the PBOC monetary frameworks have been moving towards interest rates as policy instruments. Kamber and Mohanty (2018) and Das and Song (2022) use high-frequency surprises from monetary policy announcements on the short-term repo market to measure the impact of China’s monetary policy.

In this paper, I compare the dynamic effects of monetary policy shocks computed using the M2 growth shocks in Chen et al. (2020) and the high-frequency surprises from Das and Song (2022). I apply an empirical model that allows for both credit and interest rates to have a role in the transmission of monetary policy. I also contribute to the literature by suggesting prior restrictions, such that median responses to monetary policy shocks are effectively temporary. Finally, I provide additional evidence supporting the role of China’s monetary policy in driving the global economy.

Comparing the two alternative proxies for monetary policy in China suggests that interest-based policy has delayed effects on output and prices compared with credit-based policy. The development of the domestic financial system and the increasing use of policy rates as the main policy instrument may lead to less effective control of fluctuations in output and prices in China.

I also show how applying a “temporary effect” prior improves our understanding of long-run effects estimates for persistent data, helping with the bias-variance trade-off reported in Li, Plagborg-Møller and Wolf (2022). The prior leads to impulse responses that clearly characterise the temporary nature of monetary policy shocks on output, easier to reconcile with DSGE responses.

Miranda-Agrippino and Rey (2022) report evidence that China’s monetary policy has
a role in the global trade and commodity cycle, but not in the global financial cycle. I pro-
vide evidence that a tightening of monetary policy in China leads to a significant increase
in excess bond premium, which is a measure of global financial risk (Gilchrist, Wei, Yue
and Zakrajsek, 2022). This effect peaks one year after the shock and it is associated with
a decline in international trade.

In contrast to Barcelona et al. (2022), but in line with Chen et al. (2018), I use official
China data to measure economic activity. Economic activity is measured by seasonally
adjusting a quarterly industrial production index. The dampening effect, reported in
Barcelona et al. (2022) and partially replicated in a robustness exercise in this paper, is
not evident with industrial production.

The two proxies for monetary policy in China are described in Section 2, including
an analysis their relevance as instruments. Section 3 describes the SVAR-IV approach
and the main empirical comparison of this paper. I also included empirical responses
using the alternative internal instrument approach. Results suggest that the increasing
parameter uncertainty of the alternative approach masks the effects of monetary policy
in China.

In section 4, I describe how to implement the temporary effect prior and present the
results of the application to the China monetary policy proxies comparison. Section 4
also provides evidence of how the use of GDP instead of industrial production to mea-
sure economic activity in China leads to excessive uncertainty on the estimation of the
responses in line with the dampening effects reported in Barcelona et al. (2022). Finally,
in section 5, I discuss my evidence on the international transmission of China’s mone-
tary policy, including immediate effects on commodity prices and delayed by significant
effects on the excess bond premium.

2 Measuring China’s Monetary Policy

In this section, I describe two proxies for the People’s Bank of China (PBOC) monetary
policy. This section also provides evidence of the relevance of each proxy as an instru-
ment for monetary policy shocks in an IV approach to the identification of monetary
policy.
2.1 Monetary Policy Proxies

Chen et al. (2018) argue that a Taylor rule for a policy rate does not apply to the Chinese economy because it is hard to define potential output or trend growth for a transition economy with a rising share of investment to GDP ratio. In addition, they argue that no single interest rate is a monetary target over the 1990-2005 period. The authors propose a rule for M2 growth instead, as the central government usually sets a target for M2 growth that is consistent with targets for GDP growth and inflation for each year.

To assess the role of credit-based monetary policy in China, the Chen et al. (2018) monetary rule is estimated over the 2001Q1-2019Q4 period, extending the authors’ sample period by four years.\(^1\) As GDP growth targets were set aside during the period of the Covid zero policy, I do not extend the sample beyond 2019.

Assume that \(g_{m2t}\) is quarterly growth rate of M2 monetary aggregate (\(g_{m2t} = 100(\log(M2_t) - \log(M2_{t-1}))\) where M2 is seasonally adjusted quarterly M2 in CNY), that \(g_{xt}\) is the year on year growth rate of GDP (\(g_{xt} = 100(\log(GDP_t) - \log(GDP_{t-1}))\) where GDP is seasonally adjusted quarterly GDP in CNY, and that \(\pi_t\) is quarterly inflation at annualised rate (\(\pi_t = 400(\log(P_t) - \log(P_{t-1}))\)). The monetary rule is:

\[
g_{m2t} = \gamma_0 + \gamma_x g_{gapt-1} I[g_{gapt-1} \geq 0] + \gamma_2 g_{gapt-1} I[g_{gapt-1} < 0] + \gamma_{AR} g_{m2t-1} + \gamma_\pi (\pi_{t-1} - \pi^*) + \xi_{m2},
\]

where \(g_{gapt} = g_{st} - g_{st}^*\), \(I[\cdot]\) is an indicator function, \(g_{st}^*\) is the annual target of GDP growth, and \(\pi^*\) is target inflation (which is set 3.5% as an approximation).

In contrast, Kamber and Mohanty (2018) and Das and Song (2022) use high-frequency surprises from monetary policy announcements on the short-term repo market to measure the impact of China’s monetary policy. They argue that the Chinese financial system is developing fast, and interest rates are now likely to play a leading role in the transmission of monetary policy. They agree that there are many candidates for the "policy rate" role based on how the PBOC conducts monetary policy, so they suggest using high-frequency monetary policy surprises in line with the recent literature for the Fed (for example Gertler and Karadi (2015)) and ECB (Jarocinski, 2022) monetary policies.

\(^1\)For the regression in (1), data are extracted from the Atlanta Fed database - https://www.atlantafed.org/cqer/research/china-macroeconomy – in line with original estimates.
I use the surprise data from Das and Song (2022). They use monetary announcements scrapped from the PBOC website to identify the event dates. Surprises are measured using the "daily close-to-close change in the rate on one-year interest rate swaps based on the interbank 7-day repo rate" (Das and Song, 2022, p.11). The event-dated surprises downloaded from the Wenting Song website are converted into quarterly data using the method described in Almgreen, Gallegos, Kramer and Lima (2022). The technique considers where the event was in the quarter when computing the surprise’s quarterly measure.

As argued in Bauer and Swanson (2022), market-based surprises may be contaminated by information effects leading to a correlation between the surprises and financial and economic data available at the time of the announcement. As a consequence, I remove possible information effects using the following regression:

$$s_t = \alpha_0 + \alpha_1 \pi_{t-1} + \alpha_2 g_{t-1} + \xi_t.$$  

Note that $g_{t-1}$ is the quarterly growth rate of seasonally adjusted GDP. The regression in (2) is estimated over the 2007Q1-2020Q1 period, and the $R^2$ is 18%, suggesting evidence of informational effects. These effects were not removed by Kamber and Mohanty (2018) and Das and Song (2022).

Figure 1 presents the estimates of the credit-based, $\xi_t^{m2}$, and the interest-based, $\xi_t^s$ measures of monetary policy for China. We flip $\xi_t^{m2}$, so positive values indicate a monetary policy tightening for both measures. An inspection of the figure suggests that these measures frequently categorise the stance of monetary policy differently (tight/loose). They agree over key periods of easing, such as 2008Q4-2009Q1 and 2015Q1-2015Q2, and a period of tightening (2017Q1-2017Q2). It is also clear that $\xi_t^s$ is more persistent than $\xi_t^{m2}$.

2.2 The reduced-form VAR

The monetary policy measures in Figure 1 can be seen as two different proxies of the non-systematic PBOC monetary policy. Before evaluating the role of these proxies in computing impulse responses from monetary policy shocks, I consider the relevance of these variables as proxies for monetary policy shocks. This implies assessing their strength as
an instrument in an SVAR-IV approach, as in Stock and Watson (2018). This section describes the reduced-form vector autoregressive (VAR) model to appraise whether these proxies are able to identify monetary policy shocks.

The VAR model has eight endogenous variables. There are two economic activity variables, Output, measured by a seasonally-adjusted quarterly index of industrial production, and Exports (in US$, EXP). A measure of aggregate prices, a seasonally adjusted, quarterly CPI, is also included. Four monetary variables are included in the VAR. Two are "price" measures: the seven-day repo rate in the inter-bank market and the 1-year Medium-term lending facility rate. The other two are "quantity" measures: the reserve requirement ratio (for major banks) and the nominal growth of the M2 monetary aggregate. Many monetary variables are included because of the constellation of monetary policy instruments employed by the People’s Bank of China (PBOC) during the last 15 years, as described in Jones and Bowman (2019). The last endogenous variable is the real effective exchange rate (REER) computed by the BIS.

Notes: The repo surprise is based on changes in the 1-year interest rate swaps on the 7-day repo as provided by Das and Song (2022). The values above are the residual of a regression of the surprises on lagged quarterly GDP growth and inflation. The M2 rule is the residual of the asymmetric M2 growth rule by Chen et al. (2018) estimated over the 2001-2019 period with data from the Atlanta Fed database - https://www.atlantafed.org/cqer/research/china-macroeconomy

2Quarterly industrial production is employed to measure output instead of GDP because the discussion in Barcelona et al. (2022) on the usefulness of GDP to measure economic activity in China over most of the 2010’s decade. This is also in line with the approach in Kamber and Mohanty (2018) that uses IP as a measure of output. I show results using GDP instead in section 4

3For quarters before 2016Q1, the 1-year benchmark rate is employed to capture changes in the 1-year rate.
Let $y_t$ be a $N \times 1$ vector that includes all the eight quarterly endogenous variables, with OUT, EXP, CPI and REER in log levels (*100). Then, we can write:

$$y_t = [OUT_t, CPI_t, EXP_t, yMLF_t, 7dRepo_t, RRR_t, M2gr_t, REER_t]'$$

The reduced-form VAR(p) model is:

$$y_t = c + A_1 y_{t-1} + ... + A_p y_{t-p} + u_t$$
$$u_t \sim N(0, \Sigma), t = p + 1, ..., T,$$

and the reduced-form errors are:

$$u_t = [u^OUT_t, u^CPI_t, u^EXP_t, u^yMLF_t, u^7dRepo_t, u^RRR_t, u^M2gr_t, u^{FX_t}]'.$$

Following the recommendation of Higgins, Zha and Zhong (2016) to obtain accurate forecasts for China with VAR models, the system in eq. (3) is estimated using Bayesian methods. The autoregressive order is set to 5. We use both Minnesota and the cointegration priors. This last prior is included as the variables in the model are likely to be cointegrated, even if of unknown form. We use an MCMC sampler to estimate the prior hyperparameters using the methodology in Giannone, Lenza and Primiceri (2015). The sample period employed in estimating the VAR parameters is 2007Q1 to 2022Q3.

If the monetary policy shock proxies $\xi_{m2}^t$ and $\xi_{s}^t$ have information to identify China’s monetary policy shocks, I expect them to be highly correlated with the reduced-form errors of the four monetary variables. They may still be mildly correlated with the other reduced-form shocks as they are linear combinations of the underlying unobserved structural shocks. A key identification assumption is that the proxy correlates with the unobserved structural monetary policy shock, but not with other structural shocks that may drive the macroeconomic variables. Next, I analyse the empirical evidence on proxy relevance.
Table 1: $R^2$ of regressions between reduced-form errors and MP proxies

<table>
<thead>
<tr>
<th></th>
<th>$\tilde{\xi}^s_{t}$</th>
<th></th>
<th>$\tilde{\xi}^{m2}_{t}$</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$q^{16}$</td>
<td>mean</td>
<td>$q^{84}$</td>
<td>$q^{16}$</td>
</tr>
<tr>
<td>$u_{t}^{OUT}$</td>
<td>0.00</td>
<td>0.03</td>
<td>0.07</td>
<td>0.00</td>
</tr>
<tr>
<td>$u_{t}^{EXP}$</td>
<td>0.00</td>
<td>0.01</td>
<td>0.02</td>
<td>0.00</td>
</tr>
<tr>
<td>$u_{t}^{CPI}$</td>
<td>0.00</td>
<td>0.02</td>
<td>0.04</td>
<td>0.00</td>
</tr>
<tr>
<td>$u_{t}^{repo}$</td>
<td>0.08</td>
<td>0.15</td>
<td>0.21</td>
<td>0.04</td>
</tr>
<tr>
<td>$u_{t}^{MLF}$</td>
<td>0.13</td>
<td>0.21</td>
<td>0.28</td>
<td>0.00</td>
</tr>
<tr>
<td>$u_{t}^{M2}$</td>
<td>0.00</td>
<td>0.01</td>
<td>0.02</td>
<td>0.33</td>
</tr>
<tr>
<td>$u_{t}^{RRR}$</td>
<td>0.08</td>
<td>0.14</td>
<td>0.20</td>
<td>0.00</td>
</tr>
<tr>
<td>$u_{t}^{FX}$</td>
<td>0.02</td>
<td>0.05</td>
<td>0.09</td>
<td>0.00</td>
</tr>
</tbody>
</table>

Table shows the mean and the 16% and 84% quantiles of the $R^2$ values for each combination of proxy and reduced-form error. $R^2$ is computed using the residuals for each kept posterior draw of the parameters (5,000). The sample period is 2007Q1 to 2019Q4.

2.3 Evaluating the relevance of the proxies

The sample size available for both proxies is 2007Q1 to 2019Q4. Table 1 shows the mean and the 16% and 84% quantiles of the empirical distribution of $Rs^2$. These are computed using 5,000 draws of the parameter matrices. For each parameter draw, reduced-form residuals are calculated and then regressed against either of the two proxies.

An inspection of the table suggests that the M2 growth rule proxy correlates highly with the reduced-form M2 growth shocks. The correlation with the 7-day repo reduced-form residual is more minor ($R^2$ is 9% instead of 43%), but it is rarely zero. In contrast, the high-frequency surprises are not correlated with M2 growth innovations but with reduced-form shocks to interest rates and the reserve requirement rate. $\tilde{\xi}^s_t$ also has a mild correlation with reduced-form innovations to the exchange rates. Both proxies are not correlated with reduced-form shocks to the macro variables (OUT, Exports and CPI).

We can conclude that $\tilde{\xi}^s_t$ is a reasonable proxy for interest-based monetary policy, with the highest correlation found for the 1-year medium-term lending facility rate. In contrast, $\tilde{\xi}^{m2}_t$ measures credit-based monetary policy’s correlation to M2 growth innovations. It is likely that $\tilde{\xi}^{m2}_t$ is a stronger instrument than $\tilde{\xi}^s_t$, as at its highest, the mean $R^2$ is twice
the obtainable with $z_t^s$. More robust instruments are expected to lead to more precise estimates of the dynamic effects of monetary policy shocks.

The next section assesses dynamic responses to monetary policy by comparing estimates obtained with each proxy. As the aim is to contrast the effects of both types of monetary policy, they are not combined. Still, the alternative of applying multiple instruments is feasible and could improve the precision of estimates.

3 Dynamic Causal effects to credit-based and interest-based monetary policy

3.1 A SVAR-IV Approach

For a given set of estimates of the coefficients in the $p \times 1$ matrices $A_1, \ldots, A_p$ in eq. (3), dynamic responses can be obtained using the MA$(\infty)$ representation of the VAR($p$). If the intercept is ignored, the reduced-form VAR can be written as:

$$A(L)y_t = u_t,$$

where $A(L) = 1 - A_1 L - A_2 L^2 - \ldots$. The MA$(\infty)$ representation is then:

$$y_t = [A(L)]^{-1}u_t = C(L)A_0 \varepsilon_t,$$  (5)

where $C(L)$ will lead to an infinite sum and $\varepsilon_t$ is a vector of uncorrelated structural shocks. Stock and Watson (2018) argue that if structural shocks $\varepsilon_t$ can be recovered from the reduced-form shocks $u_t$, then the structural VAR impulse response, $C(L)A_0$, reveals the true dynamic causal effects.

Assume that $a_{0,mp}$ is a $N \times 1$ vector that measures the impact effect of the structural monetary policy shock $\varepsilon_{t,mp}$ on the $N$ endogenous variables in $y_t$. The dynamic causal effects of monetary policy shocks are $C(L)a_{0,mp}$.

I obtain $a_{0,mp}$ as follows. For each draw of the parameters, the implied residuals, $\{\hat{u}_{i,t}\}_{i=p+1}^{T}$, are computed. Then they are regressed equation by equation on chosen proxy $\zeta_t$ (with an intercept). The vector $a_{0,mp}$ is then obtained by dividing the estimates by the
effect of the proxy on the 7-day repo. This is mathematically equivalent to the two-stage least squares step for the SVAR-IV approach in Stock and Watson (2018), as discussed in Li, Plagborg-Moller and Wolf (2022).

When considering the \( \xi_t \) proxy, additional restrictions are required to obtain dynamic responses aligned with the standard theoretical view on the effect of monetary policy shocks. I assume that monetary policy has no immediate effects on the macroeconomic variables (OUT, Exp, CPI). This is in line with the frequently employed recursive assumption to identify monetary policy shocks (Ramey, 2016).

3.2 A comparison of empirical impulse responses

Figure 2 shows the impulse responses of the application of the SVAR-IV approach using either the proxy for credit-based monetary policy, \( \xi_t^{m2} \), or for interest-based policy \( \xi_t^{s} \). The estimation period is the same for both proxies: 2007-2022Q3 to estimate the VAR parameters – dynamic transmission – and 2007-2019 to obtain \( a_{0,mp} \) using each proxy at a time. The monetary policy impact is scaled to a 25 bps change on the 7-day repo rate for both proxies.

An inspection of Figure 2 suggests that a 25 bps worth of monetary policy tightening leads to a decline in output and prices for both proxies, but the effect on credit (M2 growth) and the real effective exchange rate (REER) differs. For the credit-based policy, the 25 bps increase in the 7-day repo is achieved with one ppt decline on M2 growth, but in the case of the interest-based policy shock, the M2 change is tiny.

The impact on the Chinese currency is also different for both policy shocks. The credit-based policy tightening leads to an immediate appreciation of the renminbi, but the interest-based leads to short-term depreciation. Economic theory would have suggested an appreciation of the domestic currency from monetary policy tightening. The short-term depreciation from the interest-based policy is unexpected. Miranda-Aggripino and Rey (2022) report empirical evidence that monetary policy takes too long to impact the Chinese currency. Here, we find immediate effects. As the government control of the renminbi has changed in the last 20 years, sample periods may matter to measure the impact of monetary policy on the Chinese currency. They may explain why we do not find the sluggish effects reported in Miranda-Aggripino and Rey (2022). As the same sample
period is employed for both approaches in Figure 2, the fact that REER appreciates immediately to a credit-based policy but not to an interest-based policy supports the use of credit-based monetary policy shocks to measure the impact of monetary policy in China, as argued by Chen et al. (2018).

Even more interesting are the dissimilarities in timing the peak effects of the monetary policy tightening on output and prices. The peak effect on the output of the credit-based shock is three quarters after the shock at -0.76%, but ten quarters after the shock at -0.61% for the interest-based shock. The effects on prices are also delayed. The peak effect on output is after five quarters for credit-based policy (at -0.25%) but after 13 quarters for interest-based policy (at -0.24%). Even if the magnitude of the peak effects is not very different, the transmission is. As the PBOC moves from a credit-based to an interest-based policy, one should expect a slower transmission of monetary policy.

Bauer and Swanson (2022) reports that US monetary policy has peak effects at nine months. Here, the US empirical evidence aligns with the results using the M2 rule proxy, $\zeta^m_2$, instead of an interest-rate proxy based on a high-frequency surprise approach.

In summary, the alternative monetary policy proxies lead to dynamic causal effects with different implications for policymaking in China. The credit-based proxy is a better measure based on economic theory and international experience. The earlier discussion that the M2 growth is a stronger instrument is also supportive of dynamic responses computed using the M2 growth rule shocks as a better empirical representation of monetary policy in China.

3.3 Robustness: the internal instrument approach

Li, Plagborg-Moller and Wolf (2022) argue that the SVAR-IV approach may lead to biased estimates of dynamic causal effects if the monetary policy shock is not invertible based on the set of variables in the VAR. Our strategy, as described earlier, was to include four different measures of monetary policy in the VAR, including variables linked to credit (M2 growth and reserve requirements) and interest rates (7-day repo and 1-year medium-term lending facility rate). In this section, we add the proxy variables as endogenous variables in the VAR to check the robustness of our main results, as the internal instrument approach solves a possible invertibility issue. The internal instrument approach
has the major disadvantage of shortening the sample size available while increasing the required number of parameters for the estimation. This may lead to excessive parameter uncertainty as documented by Li, Plagborg-Moller and Wolf (2022).

Figure 3 presents the impulse responses using the same proxies as Figure 2 but using the internal instrument approach. Impact effects are computed using a Choleski decomposition of the variance-covariance matrix with the proxy ordered first. The additional recursive restrictions and the scaling to a 25 bps impact on the repo rate are then applied. As Figure 3 plots are in the same scale as Figure 2, it is clear how the 68% bands are, in general, wider. The implied dynamic effects are also bumpier than in Figure 2. Both changes are explained by the increased parameter uncertainty of the internal instrument approach applied to a short sample period (13 years of quarterly data).

The differences between the interest-based and the credit-based monetary policy are as in Figure 2. The peak effect on output is after two quarters for credit-based policy but after six quarters for interest-based policy. A similar gap in peak effects is found for prices. So the evidence of slower effects of interest-based policy is supported. The differences in the impact on the M2 growth and the REER are also sustained. The sample period impacts the size of the peak effects of M2 growth, as they decline when the 2020-2022 period is removed. A similar effect is not observed for an interest-based policy, as peak effects are around -0.56 for output and -0.26 for prices.

The dynamic effects of interest-based policy seem to be more robust to the change of approach. Figure 1 may help us explain why. The $\xi_s$ proxy exhibits more serial correlation $\xi_m$. The inclusion of $\xi_m$ in the VAR may disturb more VAR estimates due to the excessive number of spurious dynamic coefficients than the inclusion of $\xi_s$.

Our preferred measure of monetary policy effects in China is to apply the SVAR-IV approach with the $\xi_m$ proxy as it allows us to include the turbulent 2020-22 period to estimate dynamic effects and leads to the dynamic impacts of monetary policy that are line what is expected based on DSGE models and economic theory.
4 Temporary Monetary Policy Effects

The persistent hump-shape effect of the impact of the monetary policy shocks on the output displayed in Figure 2 may need to be revised to reconcile with the dynamic responses of DSGE models. Li, Plagborg-Moller and Wolf (2022) report the difficulties of obtaining unbiased long-run impulse responses for persistent data. Our Bayesian estimation using log-level data assumes a prior mean consistent with the trend behaviour in the data, helping to mitigate possible biases. The median response, however, is quantitatively non-zero for horizons after two years, even if the 68% confidence bands include zero.

If the researcher has a prior that monetary policy shocks have temporary effects on output, adding this information before estimation could improve our understanding of the dynamic effects. For example, typical algorithms to obtain the posterior distribution of VAR parameters include a step where draws from the conditional posterior that lead to non-stationarity (or that implied companion form matrix has eigenvalues outside the unit circle) are discarded. This truncates the posterior density based on the prior information of system stationarity. Similarly, the prior that monetary policy shocks have temporary effects on output can be imposed by discarding impulse response draws outside a narrow interval around zero.

As discussed earlier, the dynamic transmission of the monetary policy shock uses the inverted VAR coefficients and can be written as:

$$\Theta_j = C_j a_{0,mp} \text{ for } j = 0, 1, ..., h$$

(6)

The prior that the monetary policy has a temporary effect is imposed by discarding candidate the draws of $\Theta_{h^*}$ that do not satisfy magnitude restrictions. In particular, the value of $\Theta_{h^*,j}$, where $j$ identifies the responses of the output variable and $h^*$ is the horizon that the restriction is imposed, is assessed as $I[-w < \Theta_{h^*,j} < w]$, where $w$ depends on the size of the shock and $I[\cdot]$ is an indicator function. The value of $w$ is set such that at least 35% of the candidate draws of $\Theta_{h^*}$ satisfies the constraints. The candidate draw relies on the posterior parameter draws for $A_1, ..., A_p$ and the corresponding value of $a_{0,mp}$. For each time the restriction is accepted, the complete response function $(\Theta_0, ..., \Theta_h)$ is kept

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for all variables. Then we use the \( J \) kept draws to compute the median as the central estimate and the 68% credible intervals as in Figure 4.

The approach above can be applied to temporary shocks from other sources if there are candidate responses that satisfy the prior. I suggest computing the responses without the prior restriction to observe whether credible intervals include instances that satisfy the constraints.

We apply the temporary-effect prior approach to the SVAR-IV model with the two proxies described in Figure 2. Figure 4 shows the results for \( h^* = 15 \) and \( w = 0.04 \) applied to the output response.

By comparing the estimates with the temporary-effect prior (Figure 4) and without it (Figure 2), we observe that the prior has tiny effects on the dynamic responses of M2 growth, REER and the 7-day repo. This suggests that the temporary effect prior is compatible with both types of monetary policy. As expected, the median responses of output are lower than without the truncation achieved by the prior. The timing of the peak effects on output and prices is shortened using the interest-based proxy. But still, this paper’s main result –interest-based policies have a delayed impact compared with credit-based ones– is robust to imposing the temporary effect prior. The responses to interest-based policy are, however, more susceptible to the restriction at \( h_s = 15 \) as initial responses had later peak effects. Indeed the size of the peak effect on output declines from -0.61 to -0.31 and from -0.23 to -0.12 for the CPI peak. The 68% confidence bands for the dynamic responses of output and prices using the \( b_{1t} \) always include zero. In contrast, we still find evidence that effects over the first two years are negative using the \( b_{1t}^{m2} \) proxy.

These empirical results support the use of the M2 growth rule to measure monetary policy shocks in China, even if we truncate the distribution of the responses to exclude cases in which long-run effects of policy on output are very persistent. A monetary policy that cuts M2 growth in about one ppt and lets the repo rise in 25 bps leads to a decline in output of -0.70% three quarters after the policy change. The peak effect of the policy on CPI is of 0.21% four quarters after the policy change.
4.1 Robustness to the use of alternative economic activity variable

The benchmark results reported earlier use industrial production as a measure of economic activity. The main reason is that GDP is too smooth between 2012 and 2019 as reported in Barcelona et al. (2022). Preliminary evidence suggests that quarterly IP growth averages are usually in line with quarterly GDP growth averages over the 2007-2022 sample period, but IP growth is not as smooth as GDP growth. Here we evaluate the robustness of the results in Figure 4 to use GDP as a measure of activity. The recursive restrictions on output, prices and exports are applied for both proxies.

Figure 5 presents the results of the empirical exercise just described. The dynamic responses of M2 growth, REER and the 7-day repo are only superficially impacted by the use of GDP as a measure of activity in the VAR. The responses computed with the $\xi^m_t$ proxy for output and prices are shallower than before, also reflecting the effects of the imposition of the recursive restriction. The dynamic impact of the interest-rate-based policy is still delayed compared to the credit-based, but now the interest-based policy has a more substantial effect on output, particularly at the two-year horizon.

The dampening effect reported in Barcelona et al. (2022) is replicated here but only if using a credit-based measure of monetary policy.

Figure 6 replicates the same exercise as Figure 5, but using the estimate for China’s GDP from Barcelona et al. (2022). As data is only available up to 2019, the estimation period available for the VAR dynamics is shorter. The responses to credit-based policy are similar to those obtained when using GDP instead of IP, but confidence bands are tighter. Responses of output are attenuated for the interest-based policy if compared with 5. I can then conclude that the choice of using IP to measure output is not detrimental to comparing monetary policy effects in China.

5 International Transmission of China’s monetary policy

Miranda-Aggripino and Rey (2022) show with the aid of vector autoregressive models that the US Federal Reserve plays a vital role in the global financial cycle as the Fed

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$^4$Monthly GDP estimates were downloaded from Danilo Cascaldi-Garcia website, and converted to quarterly frequency by averaging over the quarter.
monetary policy affects global financial market variables. In contrast, the PBOC policy drives the worldwide trade and commodity cycle, but they have a limited effect on the global financial cycle. Barcelona et al. (2022) find that 1% demand shock (credit impulse) leads to an increase in commodity prices of 2.2% and substantially impacts global trade and global industrial production. China’s credit impulse shock does not affect the VIX supporting the claim that China has no role in the global financial cycle.

In this section, I compare the international transmission of China’s monetary policy measured using the M2 growth rule proxy, $\xi_{m2}$, with the transmission of the US monetary policy as proxied by the orthogonal surprises in Bauer and Swanson (2022). As before, the dynamic responses are estimated using an SVAR-IV approach and for the same sample period (2001-2022Q3).

Following the VAR specification in Miranda-Aggripino and Rey (2022), we estimate a quarterly VAR model for the following variables: (i) GDP and CPI in log levels; (ii) the 2-year treasury bill as a measure of the monetary policy stance as in Bauer and Swanson (2022); (iii) the US real effective exchange rate; (iv) the excess bond premium (EBP) as a measure of global financial risk as suggested by Gilchrist et al. (2022); (v) a commodity price spot index; (vi) world trade and world industrial production from the CPB-WTM as in Miranda-Aggripino and Rey (2022). The equivalent VAR for China has the same set of global variables (EBP, commodity prices, world trade and world industrial production), but instead the following domestic variables: (i) quarterly IP index and CPI in log levels; (ii) the 7-day repo and M2 growth, and (iii) real effective exchange rate.

Both VAR specifications are estimated as described in section 2.2 with p=5, and impact effects are computed as in section 3.1. The proxy for US monetary policy shock is the orthogonal surprise in Bauer and Swanson (2022) converted into quarterly sampled data from event-dated values as in Almgreen et al. (2022). I use the M2 growth rule proxy for China as justified previously.

Figure 7 shows the dynamic responses to a 25 bps monetary policy shock. For China, the 7-day repo rate is used to scale the shock size (equivalent to 1 ppt decline in M2 growth). For the US, the role is performed by the 2-year Treasury bill. The VAR to compute the responses to the US policy shock has eight variables, as previously listed. The

$^5$Many of these variables were converted to the quarterly frequency by averaging over quarter.
one for China has nine variables, as two related monetary policy variables are required (M2 growth and the repo rate). The sample period for the policy proxies for both countries is 2001-2019.

The Fed monetary policy surprise has significant and sizeable effects on world industrial production and trade. The responses to the PBOC monetary policy are generally smaller, but the world IP 68% bands do not include zeros in horizons from 2 to 6. The shape of the responses to the China monetary policy is in line with the credit impulse results in Barcelona et al. (2022). Still, the effects are smaller, likely caused by the fact that we need to build an alternative measure of China’s GDP, as argued by the authors. But compared to Barcelona et al. (2022), we find faster effects of China’s monetary policy on commodity prices.

Based on the results in Figure 7, it is clear that commodities and trade are relevant channels of the international transmission of China’s monetary policy, as argued by Miranda-Aggripino and Rey (2022). The Fed’s effects on commodity prices are muted compared with the PBOC policies. As expected, the measure of global financial risk (EBP) significantly tightens at impact as the US policy surprise hits. Interestingly, China’s monetary policy tightening also increases global financial risk, albeit with a delayed effect of about a year. The hump-shaped result is in line with the impact of China’s policy on world trade.

In summary, these results provide additional evidence of how China’s monetary policy transmits to the rest of the world. Effects are slower than the ones of the US monetary policy via the financial risk channel, but they are sizeable and mainly via the trade and commodities channel. We also provide new evidence of effects on the financial risk channel, even if not immediate, as in the case of the US monetary policy.

6 Conclusion

The empirical evidence in this paper supports asymmetries in the dynamics of the transmission of monetary policy. Depending on whether the policy is implemented using changes in the quantities of credit or interest rates, output and price responses peak at different horizons. Interest-rate-based policies have delayed peak effects compared to
credit-based ones.

The empirical results support the claim that changes in the PBOC policy implementation towards targeting a policy rate may change the dynamic of the transmission mechanism compared to the previous policy of targeting M2 growth. The evidence also supports macroeconomic models that consider monetary policy rules for China that do not rely on a single policy rate as it is usually implemented for other countries with the assumption of a Taylor rule.

References


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Figure 2: Dynamic Causal Effects of a 25 bps change in the 7-day repo rate using the M2 rule and the high-frequency surprise as proxies for the monetary policy shock

Notes: The VAR estimation period is 2007Q1-2022Q3 and proxies are available up to 2019Q4. These are based on VAR with eight variables: output, CPI, exports, REER, 7-day repo, 1-year MLF, RRR and M2 growth. The dotted lines are 68% credible intervals obtained using 15000 draws of the posterior distribution of VAR parameters.
Figure 3: Dynamic Causal Effects of a 25 bps change in the 7-day repo rate using the M2 rule and the high-frequency surprise as proxies for the monetary policy shock: Internal Instrument Approach.

Notes: The VAR estimation period is 2007Q1-2019Q4. These are based on VAR with nine variables: proxy, output, CPI, exports, REER, 7-day repo, 1-year MLF, RRR and M2 growth. The dotted lines are 68% credible intervals obtained using 15000 draws of the posterior distribution of VAR parameters.
Figure 4: Dynamic Causal Effects of a 25 bps change in the 7-day repo rate using the M2 rule and the high-frequency surprise as proxies for the monetary policy shock: with the temporary effects prior.

Notes: The VAR estimation period is 2007Q1-2022Q3, and proxies are available up to 2019Q4. These are based on VAR with eight variables: output, CPI, exports, REER, 7-day repo, 1-year MLF, RRR and M2 growth. The dotted lines are 68% credible intervals using the accepted draws (within the long-run constrain on output responses at horizon 16).
Figure 5: Dynamic Causal Effects of a 25 bps change in the 7-day repo rate using the M2 rule and the high-frequency surprise as proxies for the monetary policy shock: with the temporary effects prior and GDP as activity measure.

Notes: The VAR estimation period is 2007Q1-2022Q3, and proxies are available up to 2019Q4. These are based on VAR with eight variables: GDP, CPI, exports, REER, 7-day repo, 1-year MLF, RRR and M2 growth. The dotted lines are 68% credible intervals using the accepted draws (within the long-run constrain on output responses at horizon 16).

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Figure 6: Dynamic Causal Effects of a 25 bps change in the 7-day repo rate using the M2 rule and the high-frequency surprise as proxies for the monetary policy shock: with the temporary effects prior and BCHV-GDP as activity measure.

Notes: The VAR and proxies’ effects estimation period is 2007Q1-2019Q4. These are based on VAR with eight variables: BCHV-GDP, CPI, exports, REER, 7-day repo, 1-year MLF, RRR and M2 growth. The dotted lines are 68% credible intervals using the accepted draws (within the long-run constrain on output responses at horizon 16).
Figure 7: Dynamic Causal Effects of a 25 bps change in the 7-day repo rate (China) and the 2-year TB (US)

Notes: These are based on two different VAR models estimated over the 2001-2022Q3 sample period. The one to measure the effect of the US has US GDP, US CPI, 2-year TB rate, US REER, EBP, commodity prices, world trade and world IP. For the China impact, the VAR includes China IP, China CPI, 7-day repo, China M2 growth, China REER, EBP, commodity prices, world trade and world IP. The instrument for the China monetary policy is the M2 growth rule, and for US monetary policy is the orthogonalised surprise in Bauer and Swanson (2022). Both instruments are available up to 2019Q4. The dotted lines are 68% credible intervals obtained using 15000 draws of the posterior distribution of VAR parameters.