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Regional effects of monetary policy in China

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Abstract
Unitary monetary policy in large emerging economies with substantial regional disparities is likely to have heterogeneous effects with unintended consequences. This paper explores the regional effects of monetary policy in China by estimating the response of a series of provincial variables to a national monetary policy shock using quarterly data over the period 1999-2022. Regional heterogeneity is assessed by comparing the results from a fixed-effects and a mean-group estimator. The response of consumer prices and loans is found to be homogeneous across provinces, while that of output and property prices exhibits significant regional variation. Further analysis of the differential response for two provincial clusters indicates that output in Western China experiences faster drops after a contractionary monetary policy shock and takes longer to recover than in Eastern and Central China. In the same context, property prices react with a delay and endure a more gradual recovery after the shock. The advancement of market institutions, the share of state-owned enterprises, and the size of the private sector are identified as potential determinants of the differential response across the two regional clusters.

Keywords monetary policy · regional effects · China

JEL Classification E52 · E58

1 Introduction

Monetary authorities rely on data aggregated at the national level to formulate their policy targets in response to aggregate shocks. However, regional and local economies deviate in various aspects and to a differing degree from the national average, which increases their vulnerability to asymmetric shocks. The larger the extent of the heterogeneity, the more diverse the regional impact of a national monetary policy shock

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is likely to be, making it harder for monetary authorities to achieve their objectives. Large developing and emerging economies are particularly suceptable to this issue because they generally exhibit greater structural differences across regions due to uneven developmental strategies, industrial policies, or international trade linkages, among others.

This paper explores the regional effects of monetary policy in China, where large regional disparities prevail and are well-documented in the literature. Wang et al. (2023) show that only a few provincial business cycles coincide with the national one, while most idiosyncratic contractions at the regional level differ in timing. Moreover, energy-related regions seem more likely to have independent business cycles. Gatfaoui and Girardin (2015) also report that only coastal provinces have synchornized their cycles with the national one. Searching for an explanation, Poncet and Barthélemy (2008) point out that differences in the industrial structure and international trade intensity of Chinese provinces reduce business cycle co-movements, while inter-provincial trade and national fiscal coordination promote business cycle synchronization.

In this study, the regional effects of monetary policy in China are investigated in an empirical framework that estimates the response of a series of provincial variables to monetary policy shocks using quarterly data over the period 1999-2022. First, we identify the monetary policy shocks as deviations from a monetary policy rule, which is estimated across several different specifications. Next, we test for regional heterogeneity in the provincial response to these shocks by comparing the impulse response function (IRF) of a given provincial variable across the fixed-effects (FE) and the mean-group (MG) estimators via a Hausman test. If the null hypothesis of regional homogeneity is rejected, indicating the superiority of the MG estimator, we split the provinces into two clusters based on the similarity of their response to the shock and explore the difference between their IRFs for each variable. Lastly, we employ a probit model to explore potential determinants of regional heterogeneity.

The literature on regional monetary policy has grown rapidly since the 1970 s but it focuses almost exclusively on the developed economies in North America and Western Europe, especially in the context of the European Monetary Union (for a survey of the literature, see Rodriguez-Fuentes (2006) and Dominguez-Torres and Hierro (2019)). Recent studies on emerging economies, including Indonesia (Aginta and Someya 2022), Mexico (Torres-Preciado 2021), Pakistan (Faraz and Iftikhar 2020), and Poland (Anagnostou and Gajewski 2019), reveal a heterogeneous response of regional output, inflation, and unemployment to monetary policy shocks, pointing to industrial structure, firm size, and demography as potential sources of the cross-regional variation. Studies on China report similar findings. Cortes and Kong (2007) show that the output of coastal provinces exhibits a stronger response to shocks than for inland provinces. Guo and Masron (2014, 2017) confirm the presence of regional heterogeneity in a structural VAR framework, indicating that spillover effects across provinces play an important role in the short run. Using city-level data, Yu and Huang (2016) suggest that M2 shocks have a greater impact on housing prices in Eastern China than in the central or western parts of the country.

Our paper is closely related to recent contributions like Guo and Masron (2017) and Tsang (2024), but differs in several important respects. Guo and Masron rely on a structural VAR framework with annual data, which limits the ability to identify short-run dynamics and to flexibly accommodate regional heterogeneity over time. In contrast, we use quarterly data and a local projections approach that allows for a direct estimation of impulse response functions across multiple provincial variables. This enables us to assess not only whether monetary policy has heterogeneous effects, but also when and how such effects unfold. While Tsang (2024) employs machine learning techniques to uncover cross-sectional variation in policy effects, his theoretically largely agnostic framework focuses on average treatment effects and does not estimate dynamic responses over time. Our approach complements his by emphasizing the temporal dimension of monetary transmission and by identifying which aspects of the response are homogeneous and which show pronounced heterogeneity. Furthermore, we introduce a novel clustering procedure based on the empirical similarity of provincial impulse responses, rather than relying on predefined geographic or economic classifications.

The literature identifies bank lending as one of the leading channels of the asymmetric effect of monetary policy shocks across Chinese provinces. Cortes and Kong (2007) show that the share of loans to industrial enterprises magnifies the accumulated impulse response of provincial output, while Guo and Masron (2017) find that provinces with a large proportion of small enterprises and a small concentration of state-owned enterprises (SOEs) are more sensitive to monetary policy shocks, arguing that SOEs are less responsive because they face soft-budget constraints. Emphasizing the role of institutional factors, Liao and Wang (2021) promote the argument that inter-provincial competition and incentives for local cadres to stimulate growth are responsible for the heterogeneous response to monetary policy shocks. In particular, provincial governments facing fierce competition from neighboring regions are less concerned with national policy targets and seek to boost growth by relaxing budget constraints and extending credit to enterprises with the help of local banks, thereby weakening the effectiveness of national monetary policy.

While previous studies have established that business cycles are often asynchronous across regions, the extent to which monetary policy shocks propagate unevenly remains an open and policy-relevant question. Several recent studies have investigated this issue in specific countries, including China, but many are subject to methodological limitations that restrict the analysis of regional heterogeneity in a dynamic setting. Our study contributes to the literature and improves on exisiting research in several ways. Most of the earlier works use annual data, which severely restricts the time dimension of the sample. In contrast, one of the strengths of our paper is the exclusive use of observed quarterly data–or monthly data aggregated to the quarterly level–covering more than two decades, including the pandemic period and its aftermath. This allows for a more accurate estimation of impulse responses and a clearer assessment of regional heterogeneity in the short- and medium-term effects of monetary policy shocks. In addition, the period under investigation often begins in the late 1970 s, a time marked by a structurally different monetary regime. During China's planned economy era, the role of money was administrative rather than market-based, and monetary policy, in the modern sense, was effectively absent (Chen et al. 2018). Instead, we focus on the two most recent decades, during which monetary policy in China has become more effective as international standards were adopted and the targeting of monetary aggregates was gradually replaced with an interest-based policy framework (Kim and Chen 2022).

In terms of methodology, previous studies favor VAR models. By modeling the joint dynamics of all variables involved, VAR models are extremely helpful to build an understanding of the details of the transmission mechanism. However, this comes at a cost. The lag order of a correctly specified VAR is determined by the largest lag of the effect of any variable on any other variable in the VAR. This can lead to a large number of estimated coefficients and correspondingly high demands on the length of the available time series. For a meaningful model all variables playing a role in the transmission need to be included, intensifying the data requirements further. Structural VAR models typically need to impose restrictions to identify structural shocks properly, while in global VAR models cross-regional variation is determined by the weight matrix. By comparison, we take an agnostic approach that makes fewer assumptions about the response to a monetary shock, allowing the data to determine the effect, providing a complement to the existing literature Furthermore, we examine the IRFs of output, consumer prices, loans, and property prices, whereas existing research typically focuses on a single variable. While loans are generally treated in the liteature only as a transmission channel for asymmetric shocks, we explore loans as a variable responding to monetary policy shocks in its own right. Lastly, we cluster provinces based on the empirical similarity of their responses to monetary shocks, allowing us to move beyond stylized regional classifications and to uncover new insights into the channels and structural determinants of asymmetric monetary transmission in large, decentralized economies.

The rest of the paper is structured as follows. The next section presents the methodology, while Section 3 describes the data. Section 4 discusses the results, and Section 5 provides some conclusions.

2 Methodology

What we aim to identify is the differential reaction to monetary policy shocks in different regions of China. This requires us to first identify both the shock and the appropriate division of Chinese provinces into clusters and to verify whether there is heterogeneity at all. To this end, we first estimate a monetary policy reaction function (more specifically a Taylor-type rule) and take the deviations from that rule as monetary policy surprise (i.e., the monetary policy shock). Second, we assess whether there is a meaningful degree of provincial heterogeneity in the response of our variables of interest to those shocks. To do so, we embedd the shocks in panel local projections to estimate impulse responses to monetary policy shocks in two different models, one allowing for province level heterogeneity and one assuming homogeneity. If the data prefer the heterogeneous model for a given response variable, we move on to the third step. In this third step, we cluster province level impulse responses by similarity generating regions (clusters of provinces) with differential responses of the variable of interest to monetary policy surprises. Finally, in a fourth step we directly estimate difference in the impulse response between regions identified in step three.

At the core of our study are the key objectives of monetary policy: output and inflation. However, we also consider two additional indicators. First, we include house prices. The People's Bank of China (PBoC) itself has frequently emphasized the importance of house prices in its monetary policy reports. It has actively intervened in the housing market–initially to prevent a further buildup of a potential bubble, and more recently to manage a gradual decline in prices and avoid a sudden collapse. A large body of literature has examined the impact of the PBoC's monetary policy on real estate prices, making house prices a natural candidate for assessment in this regional context.

Second, we include loans. Loans are central to the transmission of monetary policy, not only through the credit channel–comprising both the balance sheet and bank lending channels–but also through the traditional interest rate channel. The balance sheet channel posits that the external finance premium is positively correlated with the policy rate, as higher interest rates depress asset prices and thus reduce available collateral, leading to lower credit demand under contractionary policy. The bank lending channel, on the other hand, focuses on banks' access to liquidity and how it constrains their ability to supply loans. Finally, the interest rate channel suggests that, due to financial market imperfections, higher short-term rates feed through to higher long-term rates, reducing loan demand for long-term consumption (e.g., real estate) and investment.

For an excellent summary of these channels, see the seminal work by Mishkin (1996). While loan data do not allow us to disentangle the individual transmission channels, they still offer insights into regional variation in monetary transmission. In a market with sufficient capital mobility, regional differences in loan issuance–the first stage of transmission–do not necessarily align with regional differences in spending. Analyzing loans and output separately therefore helps identify the stage at which regional variation emerges.

Identifying the monetary policy shock We begin with the identification of monetary policy shocks. In line with the seminal paper by Chen et al. (2018), among others, we interpret shocks as deviations from an estimated monetary policy rule. The corresponding empirical specification is given by:

$$mp_t = \beta_0 + \beta_1 mp_{t-1} + \beta_2 \pi_{t-1} + \beta_3 y_{t-1} + \varepsilon_t \tag{1}$$

where mp is a measure of monetary policy, π is the national inflation rate, and y is the growth rate of aggregate output. We opt for a backward-looking model due to the lack of quarterly forecasts over the entire sample period and employ OLS to estimate Eq.(1). While some papers use different methods, it has recently been shown by Carvalho et al. (2021) that (under relatively weak assumptions) even in the face of endogeneity, OLS is superior to IV for the estimation of monetary policy reaction functions. Our preferred specification uses the CHIBOR rate but we estimate policy rules for a range of different potential monetary policy measures, including the growth rate of monetary aggregates and other interest rates.

Confirming province level heterogeneity In the next step, we test for heterogeneity in the provincial response of various macroeconomic variables to the national monetary policy shocks, $\hat{\varepsilon}_t$, obtained from Eq.(1). This is accomplished by first estimating impulse response functions (IRFs) through local projections in the spirit of Jordà (2005) and then comparing these IRFs over eight quarters using an estimator that allows for regional heterogeneity to one that does not.

Note that enforcing homogeneity where it is inapproriate, yields an inconsistent estimator. At the same time, when there is no (relevant degree of) heterogeneity, allowing heterogeneity and thus estimating more coefficients creates inefficiency. We will exploit this when testing which model is appropriate.

The baseline estimator precluding regional heterogeneity is a simple fixed effects (FE) model that only allows the intercepts to differ across provinces but constrains the other coefficients to be the same. The second estimator that allows for heterogeneity is a mean group (MG) estimator, suggested by Pesaran and Smith (1995), that conducts the estimation for each province separately and reports the cross-provincial averages of coefficients. Accordingly, we compare:

$$x_{i,t+h} = \beta_1 x_{i,t-1} + \beta_2 \hat{\varepsilon}_{i,t} + \gamma_i + u_{i,t} \tag{2}$$

to

$$x_{i,t+h} = \beta_{1,i} x_{i,t-1} + \beta_{2,i} \hat{\varepsilon}_{i,t} + \gamma_i + u_{i,t}, \tag{3}$$

where x is a given provincial variable (i.e., the natural logarithm of regional GDP, CPI, house prices or loans), h is the forecast horizon, and γ_i are province fixed effects. It is worth noting that neither model includes time fixed effects, which would be preferable to control for national effects. However, the time fixed effects have to be omitted to avoid perfect multicollinearity with the policy shock. This yields vastly underestimated standard errors. We, therefore, use this model exclusively to test for heterogeneity, i.e., we ask how much residuals are reduced by allowing for heterogeneity, but abstain from economic interpretation of the results.

To determine heterogeneity in the response of Chinese provinces to national monetary policy shocks, we compare the two aforementioned estimators via a Hausman test. If the response is homogeneous, then both estimators are consistent but only the FE is efficient (null hypothesis). If provinces react differently, MG is consistent while FE is not (alternative hypothesis), since it imposes an inapproriate restriction. In other words, if the Hausman test fails to reject the null hypothesis, then the FE estimator is preferable to MG. If we reject the null hypothesis, then the MG estimator is superior to FE, which means that we need to explore regional heterogeneity in more detail.

The Hausman test is conducted separately for each provincial variable across each forecast horizon (one through eight quarters). We then employ Fisher's aggregation method to obtain the p-value for the Hausman test for each variable across all forecast horizons and ultimately across all variables.

Clustering Given that the null hypothesis is rejected, indicating the presence of regional differences, we divide provinces into groups with a similar response to national monetary policy shocks with the help of the k-means algorithm, which generates clusters by minimizing the within-cluster variance. To ensure that the subsequent clustering is not driven by a few periods with extreme results, we standardize the IRFs cross-sectionally by dividing through the cross-sectional standard deviation for every period.

Estimating differences between regional impulse response functions Differences between the IRF of the various clusters are determined by estimating a FE model for each provincial variable given by:

$$x_{i,t+h} = \beta_1 x_{i,t-1} + \sum_{k=2}^{K} \beta_k \hat{\varepsilon}_t D_k + \gamma_i + \eta_t + u_{i,t} \tag{4}$$

where x is a given provincial variable (see equations 2 and 3), h is the forecast horizon, k is a given cluster with $k = 2, ..., K$, $\hat{\varepsilon_t}$ is the national monetary policy shock in quarter t, and D_k is a dummy variable that takes the value of one, if the province is in cluster k. Besides the province fixed effects (γ_i), we also include time fixed effects (η_t), which control for any factors that are constant across Chinese provinces but vary across time. Accordingly, the national monetary policy shock is absorbed by the time fixed effects but we can still assess its impact by introducing a term into Eq.(4) that interacts the dummy variable for a given cluster with the monetary policy shock. The main coefficient of interest is β_k, which can be interpreted as the difference in the response of a given provincial variable to the national monetary policy shock between a given cluster k ($k = 2, ..., K$) and the control group represented by cluster $k = 1$. To ensure a consistent interpretation, all clusterings are rearranged, assigning the cluster that includes Shanghai the role of a benchmark. In other words, we effectively assess whether IRFs of a certain provincial cluster deviate from the control group containing China's main economic and financial center.

To have a reference point, we estimate the model:

$$\bar{x}_{t+h} = \beta_0 + \beta_1 \bar{x}_{t-1} + \beta_2 \hat{\varepsilon}_t + u_t. \tag{5}$$

where $\bar{x}$ is the mean of the given provincial variable. We prefer the provincial average specification in Eq.(5) to estimating the aggregate response with Chinese national data, because it provides a more natural comparison to provincial impulse responses, whereas national data are dominated by few large provinces.

The last part of the analysis is aimed at explaining the association of provinces with a particular cluster, which ultimately helps us determine the factors that contribute to the regional heterogeneity in the provincial response to monetary policy shocks. For this purpose, we estimate a probit model that explains whether a province belongs to the Shanghai cluster or not for a two-group clustering. Depending on the provincial variable in question, the model includes potentially relevant variables, such as trade openness, financial deepening, state ownership of industrial enterprises, the extent of marketization, and the development of factor and product markets.

3 Data

We employ quarterly data for 30 (out of a total of 31) Chinese provinces over the period 1999q4-2022q4, excluding Tibet due to the lack of data. The quarterly frequency of the observations distinguishes our approach from earlier studies that relied on annual data and required disaggregation to estimate dynamic effects. This is one of the methodological strengths of this paper, as it allows for more precise estimation of impulse responses and a clearer assessment of short- to medium-term regional dynamics. Our sample period starts in 1999 to ensure consistency with the institutional environment of modern Chinese monetary policy. Prior to the 1990s, China's financial system lacked core features of a market-based monetary regime, such as indirect instruments and a functioning transmission mechanism (Chen et al. 2018).

For the estimation of the monetary shock, we obtain national statistics on the monetary aggregate M2, three short-run (3-month) interest rates (China Interbank Offered Rate (CHIBOR), Shanghai Interbank Offered Rate (SHIBOR), and China Interbank Bond Collateral Repo Rate), CPI, and real GDP. We estimate the impulse response functions of four provincial variables, expressed as quarter-over-quarter growth rates. Monthly CPI and property prices (average square-meter price of residential real estate) cover the entire sample period, while quarterly real GDP is available from 2000q2 onwards and monthly loans (total social financing) from 2004q2 onwards. All variables are collected from the CEIC database and seasonally adjusted using the X-13 ARIMA procedure. Monthly data are transformed into quarterly by taking the corresponding three-month average.

4 Results

4.1 Monetary policy reaction function

The estimation of the monetary policy reaction function in Eq. (1) requires us to specify the monetary policy instrument, which is not a straightforward task because the People's Bank of China (PBoC) uses multiple tools, ranging from open market operations and required reserves to central bank lending and administrative measures. The reason for this institutional setup is that the monetary policy framework has evolved over the past three decades from a centrally administered system that relied on credit plans and quantity-based targets like broad money (M2) growth to a more flexible structure that introduced market-based instruments and gradually liberalized interest rates. Over the sample period, the Chinese monetary policy framework was in this transitional phase where quantity-based measures still matter but new standards similar to those in developed economies were emerging.

Existing research addresses the issue of the coexistence of quantity- and price-based monetary tools in China in various ways. Some studies estimate a standard Taylor rule, focusing on short-run interest rates like CHIBOR (Xie and Luo 2002; Zheng et al. 2012). A second group opts for M2 growth as the response variable, arguing that, especially in earlier periods, interest rates did not play an important role due to the underdeveloped financial markets (Chen et al. 2018). A third group adopts a hybrid

Table 1 Monetary policy reaction functions

	$M2$ growth	Shibor	Chibor	Repo
Constant	1.575***	0.090	0.571**	0.155
	(0.342)	(0.199)	(0.242)	(0.181)
mpi_{t-1}	0.572***	0.850***	0.749***	0.845***
	(0.088)	(0.053)	(0.055)	(0.049)
y_{t-1}	−0.030	0.090***	0.057**	0.072***
	(0.052)	(0.024)	(0.029)	(0.025)
π_{t-1}	−0.122	0.372***	0.373***	0.336***
	(0.171)	(0.098)	(0.095)	(0.084)
Obs.	94	64	94	94
R^2	0.327	0.842	0.690	0.807
$adj R^2$	0.304	0.834	0.680	0.800
AIC	279.32	92.28	169.31	140.25

Reported coefficients are estimates from the model in Eq.(1). Standard deviation in parenthesis. *** $p < .01$; ** $p < .05$; * $p < .10$

approach, pointing out that neither the Taylor rule nor the McCallum rule provide an accurate description of Chinese monetary policy because until recently there was no short-term interest rate to anchor price expectations and M2 growth is an intermediate policy target rather than an instrument under the control of the PBoC. Accordingly, Girardin et al. (2017) devise for their empirical model a composite index that combines price-based, quantity-based, and administrative tools of monetary policy in China, while Mehrotra and Sanchez-Fung (2010) and Nuutilainen (2015) estimate McCallum-Taylor hybrid models following Hall and Mankiw (1994). Kamber and Mohanty (2018) and Das and Song (2022) abandon the reaction function framework altogether and obtain monetary policy shocks for China directly from the daily movements of the 7-day repo rate around policy announcements.

Given the diversity of monetary policy instruments and intermediate targets in China, we estimate four different specifications of Eq.(1) that involve quantity-based (M2 growth) and price-based (interest rates) measures. The results in Table 1 suggest that over the sample period the growth of broad money does not respond significantly to lagged changes in output and inflation, dismissing this model as an appropriate way for identifying monetary policy shocks in China. By contrast, all three specifications involving interest rates exhibit positive and significant coefficients for the lagged output growth and inflation and the magnitudes are relatively similar, especially for inflation. These findings are broadly in line with recent changes in China's monetary policy framework that de-emphasize M2 growth as an intermediary target and increasingly promote interest rates as instruments and targets of monetary policy (Harjes 2017; People's Bank of China 2020).

Our preferred specification in Table 1 uses CHIBOR as response variable, mainly because as the first market-based interest rate indicator in China it has data availability over the entire sample period and has been employed as a proxy for monetary policy in the literature (Ren et al. 2020; Xie and Luo 2002; Zheng et al. 2012). Moreover, the

Table 2 Correlations between monetary policy shocks across specifications

	M2 growth	SHIBOR	CHIBOR	Repo	Chen et al. (2018)
M2 growth	1.00				
SHIBOR	−0.30	1.00			
CHIBOR	−0.28	0.83	1.00		
Repo	−0.33	0.82	0.55	1.00	
Chen et al. (2018)	0.74	−0.35	−0.24	−0.38	1.00

Note: Correlation coefficients between monetary policy shocks obtained from the estimation of Eq.(1) with different monetary policy instruments

Table 3 Results of the Hausman test comparing FE and MG estimator

Variable	p-value	Variable	p-value
Output	0.004***	Loans	0.999
CPI	0.999	Property prices	0.000***

The Hausman test compares the FE and MG estimators of the IRFs for the four variables in response to a monetary shock (i.e., a one percentage point increase in CHIBOR). The p-values of the Hausman test statistic are aggregated with the Fisher's method across the eight forecast horizons. *** $p < .01$; ** $p < .05$; * $p < .10$

long-term coefficient for inflation in the CHIBOR specification is 1.49, which satisfies the Taylor principle. SHIBOR is more in line with modern standards as it is based on a quote mechanism but it is reported only since 2007. For the rest of the paper, we define the monetary shock as a one percentage point increase in CHIBOR.

We further justify our choice of CHIBOR by examining the correlation between monetary policy shocks obtained from the different models. The matrix in Table 2 indicates that the CHIBOR shocks are highly correlated with SHIBOR (0.82) and repo (0.55) shocks. In addition, we compare our shocks across all models with those from the seminal study by Chen et al. (2018) who use the growth of broad money as their monetary policy instrument. The strong correlation (0.74) regarding M2 growth and the moderate negative correlations between M2 growth and interest rate shocks lend further support to the validity of our empirical specifications.

4.2 Regional IRFs

Our next objective is to check for potential heterogeneity in the provincial response to the monetary policy shocks identified in the previous section. For this purpose, the IRFs for provincial output, consumer prices, loans, and property prices are compared across the FE and MG estimators with the help of a Hausman test. The corresponding p-values of the test statistic for each forecast horizon are aggregated using Fisher's method and presented in Table 3.

It is evident that for output and property prices the null hypothesis of regional homogeneity is rejected, indicating the superiority of the MG estimator. For consumer prices and loans, the FE estimator is the preferred choice, meaning that the coefficient measuring the effect of the monetary policy shock is identical across provinces (see Eq.(2)).

For output, our results are consistent with previous studies showing that provincial output responds differently to monetary policy shocks (Cortes and Kong 2007; Guo and Masron 2017). Similar heterogeneity has been well documented in other countries, beginning with the seminal study on US regions by Carlino and DeFina (1998). Comparable findings have been reported for the euro area by Boeckx et al. (2018) and Burriel and Galesi (2018). In the Chinese context, Yu and Huang (2016) also document regional differences in the response of property prices across cities.

By contrast, the homogeneity we observe in loan responses is more surprising. Since the effect of monetary policy on GDP is largely transmitted through credit, one would expect greater variation. The interest rate channel relies on the pass-through from central bank rates to lending rates (and ultimately credit demand), the narrow credit channel reflects banks' capacity to lend, and the broad credit or balance sheet channel captures how collateral values (affected by asset prices) influence borrowing capacity.

In other large, open economies, regional differences in these channels have been shown to contribute to heterogeneous output responses. For instance, Owyang and Wall (2009) confirm regional heterogeneity in both the interest rate and narrow credit channels in the U.S.; Ciccarelli and Rebucci (2002) show the importance of the credit channel in Europe during periods of fiscal stress; and Horst and Neyer (2020) document country-level variation in loan responses to monetary policy across the euro area. For further examples, see the survey by Dominguez-Torres and Hierro (2019).

In China, the relative uniformity of loan responses may be partly explained by its highly integrated financial system. The four largest commercial banks are state-owned and operate nationwide, which reduces regional variation in policy transmission as their behavior tends to align with national monetary policy objectives (Guo and Masron 2017). This suggests that regional heterogeneity in China likely arises at later stages of the transmission process–particularly in the use of credit rather than its availability. This interpretation is supported by El-Shagi and Jiang (2020), who find that PBoC policies aimed at expanding credit in poorer regions do succeed in raising loan volumes, but fail to stimulate local GDP growth. An alternative explanation for our findings could be that regional differences in equity financing play a larger role than loan access. However, since the asset price and balance sheet channels are fundamentally similar, this seems less plausible.

We proceed with the investigation of regional heterogeneity by dividing provinces into clusters using the k-means algorithm and then comparing the IRFs across clusters for each of the four variables. After experimenting with various numbers of clusters, we determine that creating more than two clusters does not bring additional benefits as IRFs are very similar.[1] On the map of China in Figure 1, the benchmark group identified by a lighter (blue) color includes mostly provinces in Eastern and Central China. The darker (black) color is reserved for the second provincial cluster, which consists predominantly of provinces in the southwest, northwest, and northeast of the country. Although the second cluster generally exhibits lower levels of per-capita

[1] The results for three and four clusters are available from the authors upon request.

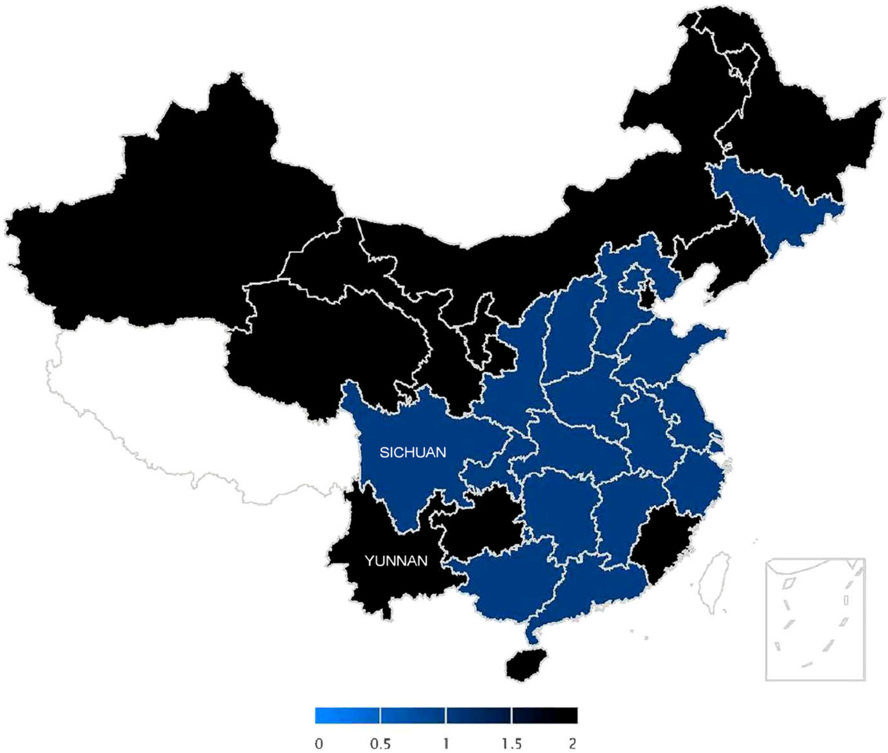


Fig. 1 The two provincial clusters of China, 1999q4-2022q4. *Note:* The benchmark provincial cluster containing Shanghai is displayed in blue and the second cluster is in black. Tibet (in white) is excluded from the sample. Clustering was achieved via the k-means algorithm, comparing the IRFs across clusters for each of the four variables in response to a monetary shock (i.e., one percentage point increase in CHIBOR). The representative provinces for each cluster (i.e., those closest to the respective cluster mean) are labeled by name on the map

income, there are exceptions, such as the coastal provinces of Liaoning, Tianjin, Fujian, and Hainan, which seem to be more similar to the poorer landlocked regions in terms of their response to a monetary shock.[2] Summary statistics for both groups, based on a range of indicators describing their economic structure, are reported in Table 4 (see Section 4.4 for a detailed description of these indicators). The table also identifies two representative provinces for each cluster–those closest to the group means based on Euclidean distance.

Figure 2 shows the impulse response of the provincial mean of the four variables to a one percentage point increase in CHIBOR. Our purpose is to ensure first that the variables respond in an economically plausible way. The contractionary monetary shock leads to a drop in all four variables as expected. Output declines for three consecutive quarters following the shock and experiences another drop in the 6th

[2] One common feature of these provinces is that although relatively rich and developed, they are overshadowed by their wealthier and more economically dynamic neighbors. For instance, Fujian is located between the economic powerhouses of Guangdong and Zhejiang, while Tianjin is adjacent to the capital Beijing.

Table 4 Summary statistics for the two clusters and their respective representative provinces

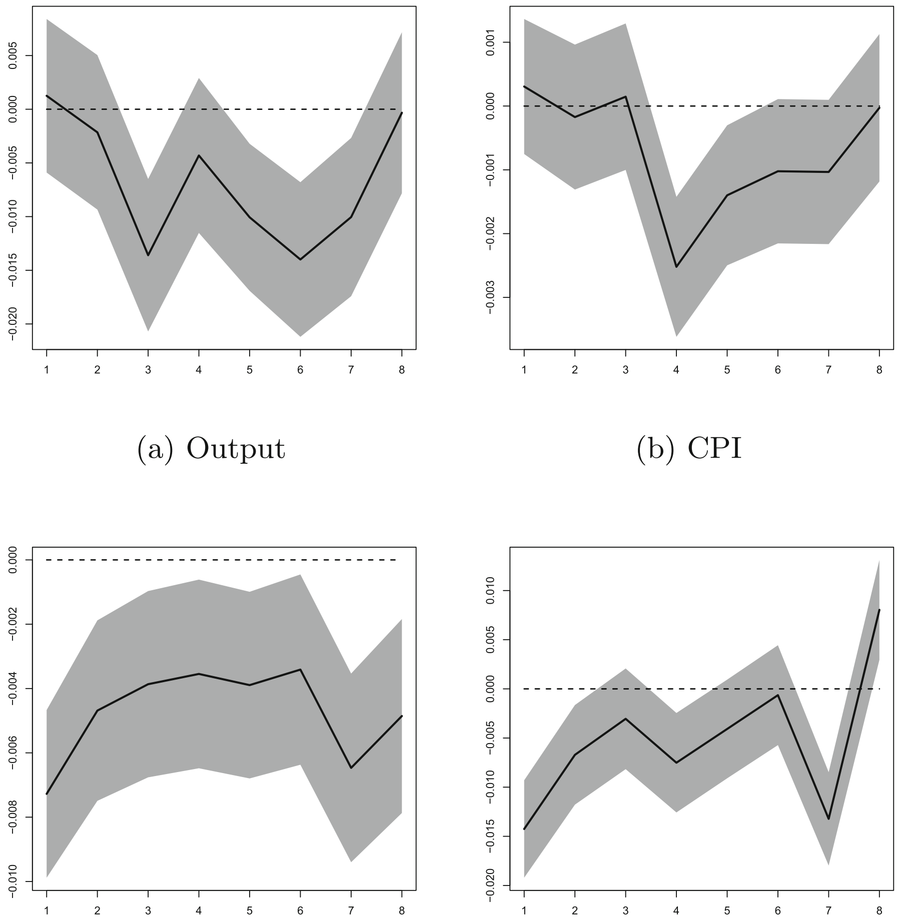
	Cluster 1 mean	Sichuan	Cluster 2 mean	Yunnan
SOE share	0.14	0.12	0.23	0.26
Foreign firms	0.07	0.03	0.06	0.03
Industry share	0.37	0.34	0.33	0.30
Fin. deepening	121.39	114.09	141.53	128.10
Big banks	13.09	12.01	16.24	13.50
Trade	34.33	11.47	24.25	9.20
Marketization	7.52	7.30	5.93	5.34
Market expansion	7.65	7.59	6.21	6.74
Factor markets	7.35	6.58	5.76	5.11
Private sector	8.76	8.59	6.82	5.91
Product markets	8.17	8.20	7.12	6.38

Representative provinces are selected as those with the smallest Euclidean distance to the cluster mean, based on indicators normalized to a mean of 0 and a standard deviation of 1. For a description of the variables, see Section 4.4

quarter before recovering. CPI records a sharp drop only in the 4th quarter, followed by a gradual recovery. By contrast, loans and property prices drop instantaneously and significantly, while the subsequent recovery is interrupted by another dip in the 7th quarter. In particular the results for GDP are quantitatively relatively large with GDP dropping by about 1 percent (a log change of slightly more than 0.01) in response to a one percentage point interest rate change increase.

Our main results are presented in Figure 3. The IRFs of output and property prices (which were found to have a heterogeneous response across provinces in Table 3) are estimated from Eq.(4) and represent the response of the second cluster to a one percentage point increase in CHIBOR *relative* to the response of the benchmark group. Output drops instantaneously, suggesting that the second cluster experiences a more severe contractionary effect than the benchmark immediately after the shock. But after a quick recovery the difference between the two clusters turns insignificant. In the 7th quarter, output declines again significantly relative to the benchmark before recovering in the 8th quarter.

As we know from Figure 2, property prices decrease instantaneously but in Figure 3 the difference between the clusters is insignificant until the 3rd quarter when the second cluster records a deep drop relative to the benchmark. However, in subsequent quarters the difference reverses, turning positive and reaching a highpoint in the 6th quarter. This suggests that property prices in the second cluster initially recover at a significantly slower pace than the benchmark but then catch up with the Shanghai cluster and even overshoot.



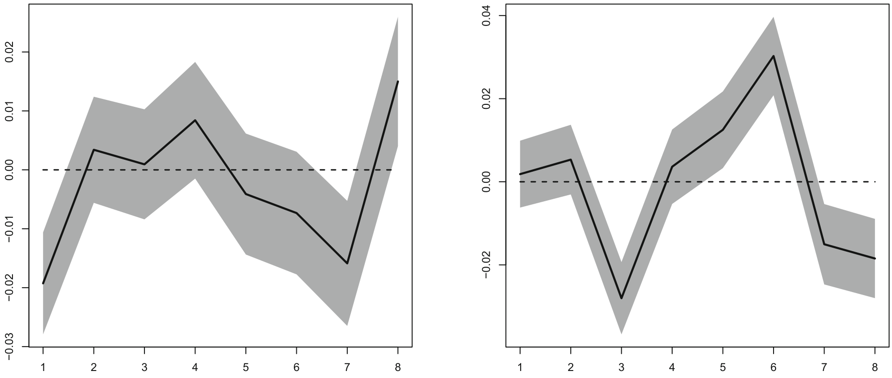
(a) Output (b) CPI

(c) Loans (d) Property prices

Fig. 2 Impulse response functions of the provincial mean. *Note:* The IRFs (with 95% confidence bounds) represent the mean response to a monetary shock (i.e., a one percentage point increase in CHIBOR) across all provinces in the sample over a forecast horizon of 8 quarters

4.3 Robustness

Our sample period includes the COVID-19 pandemic, which caused severe disruptions to the Chinese economy. We test the robustness of our results by excluding the pandemic years and limiting our sample period to 1999q4-2019q4. The map in Figure 4 suggests that the clustering has remain broadly the same. Tianjin, Liaoning, and Hainan shift to the benchmark group, while Fujian continues to be part of the second cluster. In other words, the differentiation between the coastal/central vs northwestern/northeastern provinces is more clear cut in the pre-pandemic period.



(a) Output (b) Property prices

Fig. 3 Impulse response function of second cluster relative to the benchmark. *Note:* The IRFs (with 95% confidence bounds) represent the response to a monetary shock (i.e., a one percentage point increase in CHIBOR) of the second cluster (the darker (black) color in Fig. 1) relative to the benchmark Shanghai cluster (the lighter (blue) color in Fig. 1) over a forecast horizon of 8 quarters

The IRFs of the four variables in Figure 5 are almost identical to the ones in Figure 2. More importantly, the IRF differences for output and property prices in Figure 6 are very similar to those in Figure 3, confirming the robustness of our findings. One detectable difference is that the output of the second cluster recovers somewhat faster relative to the benchmark in the pre-pandemic period than over the entire sample period.

As described in the data section, our time series begin in different years for some of the variables. For instance, monthly statistics on loans by province are available only from 2004 onwards. As a second robustness test, we conduct the analysis over the period 2004q2-2022q4, making sure that all variables are observed over the same sample period. It is evident from the map in Figure 7 that cluster membership remains largely the same. Tianjin and Liaoning join the benchmark cluster, while Shanxi switches to the second cluster. The patterns of the IRFs for the provincial means in Figure 8 and the IRFs of the second cluster relative to the benchmark in Figure 9 remain consistent with our major findings.

4.4 Determinants of regional clusters

Since we reveal regional heterogeneity in our results, we seek to identify the factors that determine the provincial composition of the two clusters, thereby explaining the variation in the regional response to monetary policy shocks. For this purpose, we employ a univariate probit regression where the dependent variable takes the value of one, if the province is part of the second cluster, and zero if it is in the benchmark cluster. The relevant statistics for the independent variables are collected from the

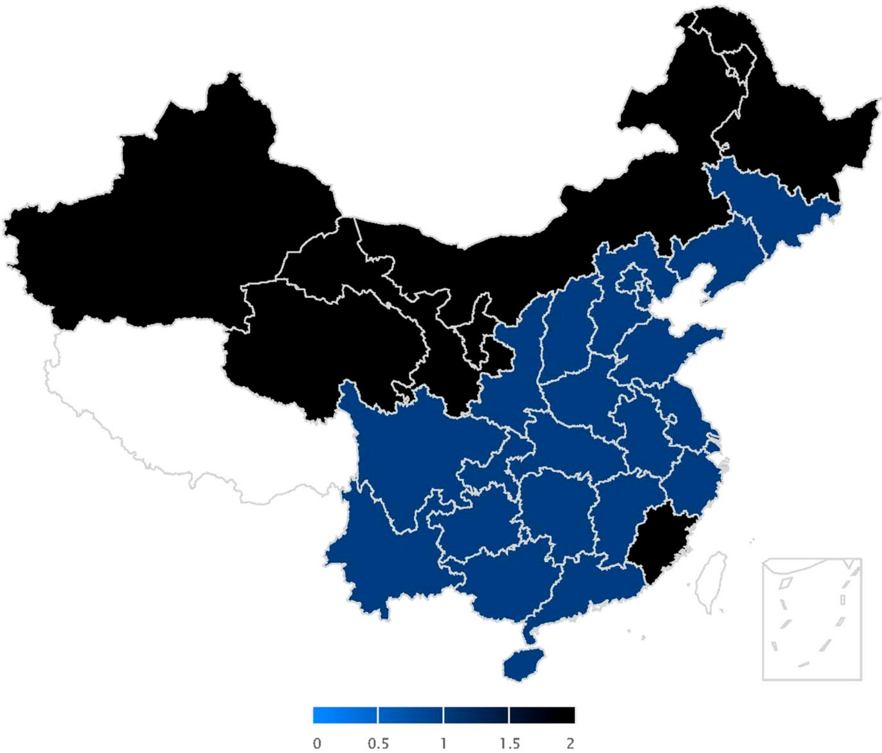


Fig. 4 The two provincial clusters in the pre-pandemic period, 1999q4-2019q4. *Note:* The benchmark provincial cluster containing Shanghai is displayed in blue and the second cluster is in black. Tibet (in white) is excluded from the sample. Clustering was achieved via the k-means algorithm, comparing the IRFs across clusters for each of the four variables in response to a monetary shock (i.e., one percentage point increase in CHIBOR)

CEIC database and the annual values for each province are averaged over the sample period.

In general, we expect that factors linked to the private sector, international integration, financial development, and the advancement of the market economy will facilitate the transmission of a monetary shock. Privately-owned companies are likely to be more responsive to market signals in an environment where government support is targeted towards the state sector. Global linkages through trade and investment typically make regional economies more competitive and thus more sensitive to changes in financial conditions. Deeper financial markets not only speed up the transmission of monetary policy measures but also ensure that the impact of such measures is more expansive. Lastly, insitututional changes that broaden the scope of market forces in the regional economy are generally more conducive for the effects of a national monetary shock because of stronger competition, market-determined prices, wages, and interest rates, and lower government hurdles.

Table 5 presents the results for various potential determinants. The coefficient for the share of SOEs in total enterprises is positive and significant, while the one for

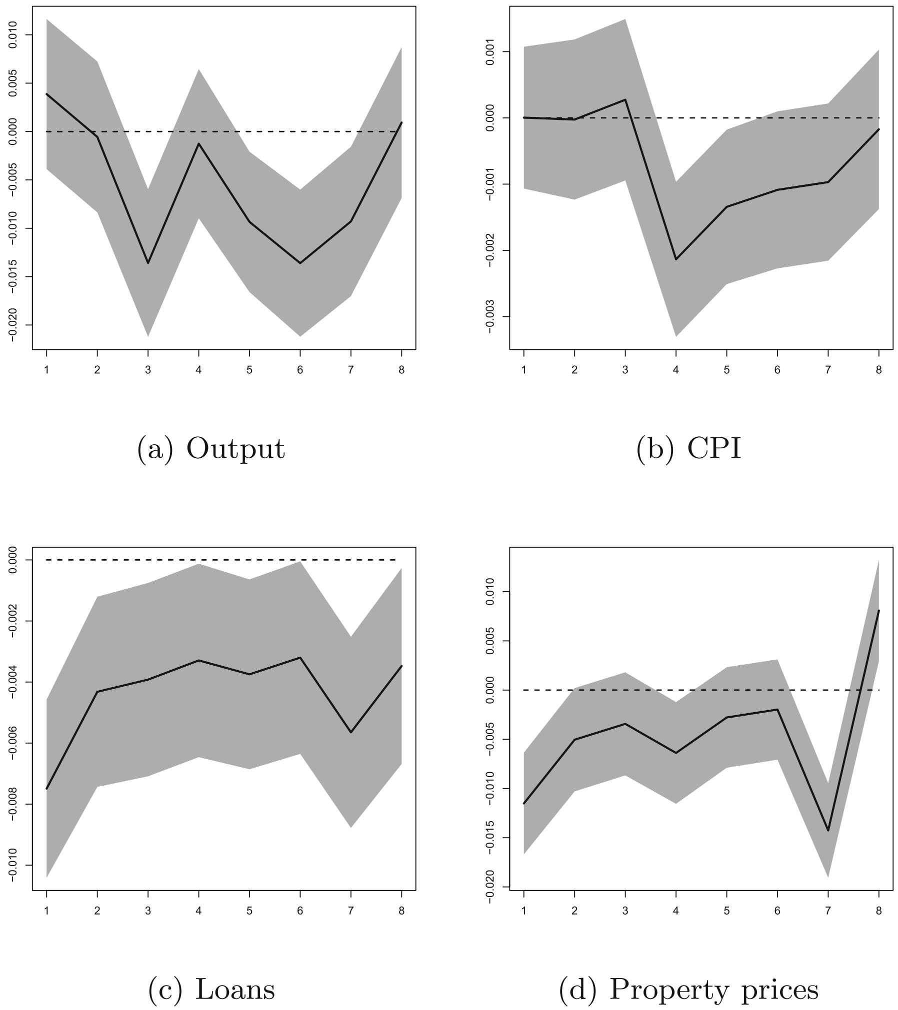


Fig. 5 Impulse response functions for all provinces, pre-pandemic period. *Note:* The IRFs (with 95% confidence bounds) represent the mean response to a monetary shock (i.e., a one percentage point increase in CHIBOR) across all provinces in the sample over a forecast horizon of 8 quarters

the share of foreign enterprises is negative but insignificant. Previous studies show that SOEs in China decrease the sensitivity of the provincial response to monetary policy shocks because these enterprises face a soft-budget constraint (Cortes and Kong 2007; Guo and Masron 2017). By contrast, foreign enterprises in general cannot rely on comparable financial support from the local governments in China. Our results indicate that the second cluster has more SOEs and fewer foreign enterprises than the benchmark, which is sensible given that foreign investment is concentrated in Eastern and Central China.

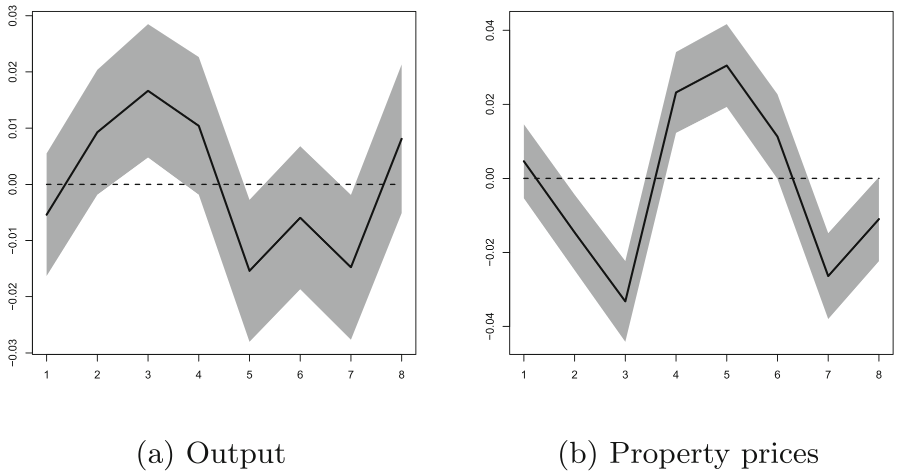


Fig. 6 Impulse response function of second cluster relative to the benchmark, pre-pandemic period. *Note:* The IRFs (with 95% confidence bounds) represent the response to a monetary shock (i.e., a one percentage point increase in CHIBOR) of the second cluster (the darker (black) color in Fig. 1) relative to the benchmark Shanghai cluster (the lighter (blue) color in Fig. 1) over a forecast horizon of 8 quarters

The GDP shares of industry and trade (exports plus imports) have a negative (albeit insignificant) impact on membership in the second cluster. The majority of provinces in that cluster are landlocked and remote, making it more difficult for them to benefit from international linkages. Regarding the importance of industry, the second cluster is rather diverse, which might be causing the lack of statistical significance. Provinces like Heilongjiang have a large secondary sector, while others like Qinghai have one of the smallest in China.

Financial deepening defined as the share of total loans in GDP has a positive and significant impact, indicating that provinces in the second cluster benefit from larger amounts of credit. This effect seems to be driven by poor Western provinces, like Gansu, Ningxia, and Qinghai, which can probably be explained by the Great Western Development Program launched by the Chinese government in 2000 and the focus on poverty alleviation and the development of the hinterlands since 2013. These government programs have provided generous funding and have made it easier to obtain loans in less developed regions concentrated mostly in Western China (El-Shagi and Jiang 2020). At the same time, the reliance on such funding has made the second cluster more sensitive to national monetary shocks, which might be responsible for the more severe contractionary effect in Fig. 3. Additional evidence comes from the GDP share of loans provided by the four largest commercial banks. The corresponding coefficient is positive and significant, supporting the argument that the provinces in the second cluster rely more heavily on loans from large state-owned banks. These banks, which are more closely aligned with central government policy, may transmit national monetary shocks more uniformly and directly than smaller or more autonomous financial institutions (Guo and Masron 2017). This institutional structure may increase the sensitivity of credit-dependent provinces to national policy shifts. While our data do not allow us to identify this mechanism causally, the interpretation is consistent with

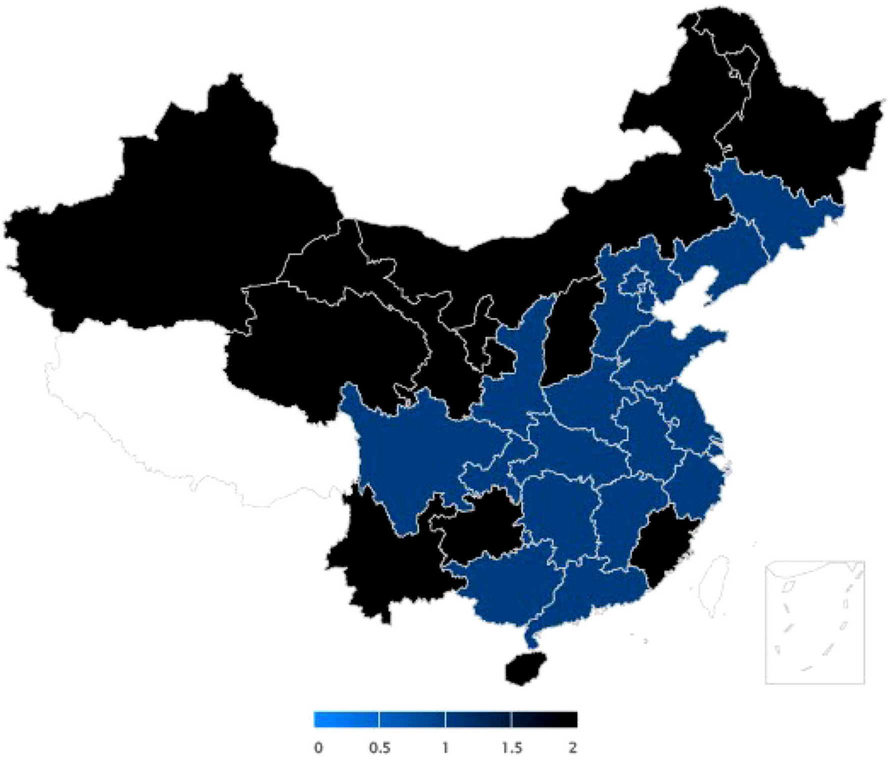
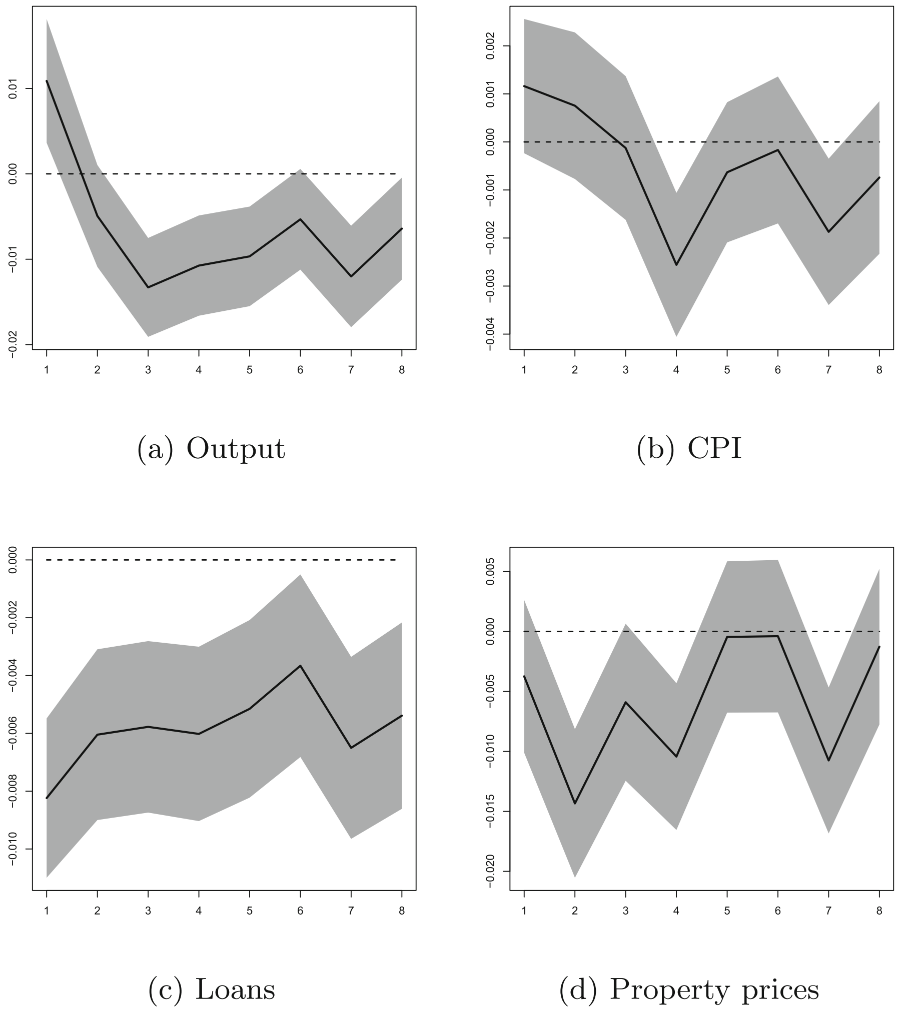


Fig. 7 The two provincial clusters, 2004q2-2022q4. *Note:* The benchmark provincial cluster containing Shanghai is displayed in blue and the second cluster is in black. Tibet (in white) is excluded from the sample. Clustering was achieved via the k-means algorithm, comparing the IRFs across clusters for each of the four variables in response to a monetary shock (i.e., one percentage point increase in CHIBOR)

prior findings on policy-driven credit allocation in China's underdeveloped regions (El-Shagi and Jiang 2020).

We also explore the role of institutional factors proxied by the NERI (National Economic Research Institute) Marketization Index, which measures the advancement of the market economy in each province (Fan and Wang 2001; Wang et al. 2021). The overall marketization index as well as its four sub-indices exhibit significantly negative coefficients, decreasing the probability of membership in the second cluster. Market expansion focuses on lowering the tax burden on enterprises, reducing bureaucratic procedures and red tape, and scaling down the government apparatus. The advancement of the factor markets looks at the mobility of labor, access to foreign capital, and competitiveness in the banking sector. The development of product markets explores the extent to which prices are set by the market and the magnitude of regional trade barriers. The development of the private sector is assessed via the share of privately-owned industrial enterprises, and non-government fixed investment and employment. The negative effect for the private sector is consistent with the positive effect for the share of SOEs.



(a) Output

(b) CPI

(c) Loans

(d) Property prices

Fig. 8 Impulse response functions for all provinces, 2004-2022. *Note:* The IRFs (with 95% confidence bounds) represent the mean response to a monetary shock (i.e., a one percentage point increase in CHIBOR) across all provinces in the sample over a forecast horizon of 8 quarters

In summary, the provinces in the second cluster are marked by a less developed market economy with a large state sector and a small private sector. Bureaucratic procedures and a heavy tax burden on enterprises, less access to foreign capital and higher trade barriers further characterize these provinces. At the same time, reliance on funding for poverty alleviation policies by the central government and a larger share of loans from the major state-owned banks in the country makes the second cluster more sensitive to national monetary shocks.

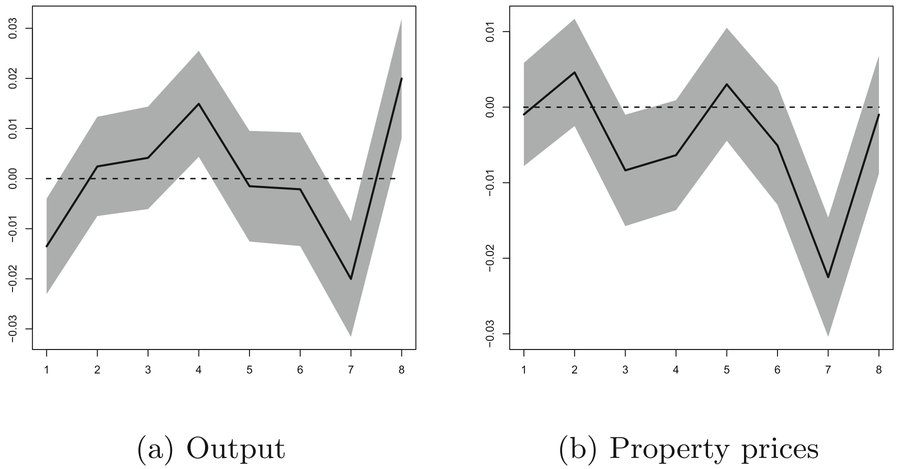


Fig. 9 Impulse response function of second cluster relative to the benchmark, 2004-2022. *Note:* The IRFs (with 95% confidence bounds) represent the response to a monetary shock (i.e., a one percentage point increase in CHIBOR) of the second cluster (the darker (black) color in Fig. 1) relative to the benchmark Shanghai cluster (the lighter (blue) color in Fig. 1) over a forecast horizon of 8 quarters

5 Conclusion

Monetary policy in large emerging economies with substantial regional disparities is likely to have heterogeneous effects with unintended consequences. We examine this issue in China, contributing to the exisiting literature by using quarterly data over the period 1999-2022, evaluating the response of a larger number of macroeconomic variables, and adopting a more flexible estimation approach that imposes less restrictions, allowing the data to determine the effect. Unlike prior studies such as Guo and Masron (2017) and Tsang (2024), which focus on average effects or rely on annual data, our approach reveals not only the existence of heterogeneity, but also its dynamic structure and institutional underpinnings, with implications for the design of regionally responsive policy tools.

Our study contributes to the literature not only by confirming the presence of regional heterogeneity in monetary policy transmission, but by advancing the analysis in several important ways. We use a long panel of true quarterly data–rather than interpolated annual series–which allows us to identify short-run dynamics with greater precision. We also apply a flexible local projections approach to estimate regional impulse responses across multiple dimensions, including output, prices, loans, and property markets. Unlike earlier studies that focus on a single variable or rely on structural VARs, we uncover that heterogeneity is primarily found in the response of output and property prices, while the response of loans is largely homogeneous. Furthermore, we systematically link these differences to structural provincial characteristics, offering new evidence on the institutional and financial roots of uneven monetary transmission.

Empirically, we estimate monetary policy shocks from a Taylor-type rule and trace their effects on provincial macroeconomic variables using both fixed-effects and mean-

Table 5 Results of the univariate probit regression

SOE share	8.744***
	(3.248)
Foreign firms	−1.825
	(3.911)
Industry share	−4.192
	(3.428)
Fin. deepening	0.013*
	(0.008)
Big banks	0.097*
	(0.057)
Trade	−0.007
	(0.008)
Marketization	−0.437**
	(0.176)
Market expansion	−0.564***
	(0.220)
Factor markets	−0.229*
	(0.122)
Private sector	−0.318**
	(0.134)
Product markets	−0.428**
	(0.211)
Obs.	29

Reported coefficients from univariate probit regression with dependent variable taking the value of 1 if a province is in the second cluster. Standard errors in parentheses. Chongqing is excluded from the sample due to missing data. *** $p < .01$; ** $p < .05$; * $p < .10$

group estimators. Where the data indicate regional heterogeneity, we cluster provinces and compare the trajectories of key variables. Rather than relying on predefined or administrative regional classifications (e.g., "coastal vs. interior"), we implement a data-driven clustering procedure that groups provinces based on the empirical similarity of their impulse response functions. We find that output in the second cluster of provinces declines more sharply and exhibits greater volatility, while property prices recover more slowly following a contractionary shock. These effects are robust to excluding the COVID-19 period and to aligning variable coverage across years. Structural factors such as the dominance of SOEs, limited marketization, and greater dependence on policy-driven credit help explain the composition of these regional clusters.

The results of our study have direct implications for the design and implementation of monetary policy in large, structurally diverse economies like China. While the PBoC's centralized framework ensures a uniform transmission of credit, our findings suggest that output and property prices respond asymmetrically across regions–particularly between the more marketized eastern provinces and the structurally less developed western and northeastern regions. To enhance policy effectiveness and avoid unintended regional disparities, we recommend that the PBoC incorporate regional economic indicators–especially for output and property markets–into its policy assessment framework. In addition, targeted credit measures or region-specific macroprudential tools could help offset the more persistent downturns observed in certain clusters. Over the longer term, efforts to deepen financial markets and reduce institutional asymmetries–such as the dominance of SOEs and underdeveloped private sectors–would contribute to a more balanced transmission of monetary policy across regions. These measures could also be coordinated with regional fiscal interventions to mitigate uneven recovery paths and foster greater synchronization of regional business cycles.

Appendix A: Methodology

Table A.1 Hausman tests for the GDP effect

horizon	FE		PMG		Hausman
	AR(1)	shock	AR(1)	shock	
1	−0.0675***	0.0015	−0.0255	0.0015	1.992
	(0.0195)	(0.0042)	(0.0351)	(0.0024)	(0.369)
2	−0.0041	−0.0020	0.0224	−0.0034	0.671
	(0.0196)	(0.0043)	(0.0340)	(0.0027)	(0.715)
3	0.0555***	−0.0134***	0.0111	−0.0135***	2.862
	(0.0197)	(0.0043)	(0.0327)	(0.0026)	(0.239)
4	−0.0992***	−0.0044	−0.0132	−0.0036	11.836***
	(0.0197)	(0.0043)	(0.0317)	(0.0023)	(0.003)
5	0.1067***	−0.0098**	0.0889***	−0.0101***	1.581
	(0.0198)	(0.0044)	(0.0243)	(0.0026)	(0.454)
6	−0.0257	−0.0139***	0.0066	−0.0139***	4.711*
	(0.0200)	(0.0044)	(0.0130)	(0.0030)	(0.095)
7	0.0085	−0.0102**	−0.0036	−0.0099***	3.551
	(0.0202)	(0.0044)	(0.0170)	(0.0036)	(0.169)
8	0.01073	−0.00036	0.03822*	−0.00068	7.821**
	(0.02029)	(0.00447)	(0.02009)	(0.00397)	(0.02)
Fisher-test					35.024***
					(0.004)

Note: The table compares the FE model to the PMG model. While the PMG model is estimated allowing for heterogeneity in slope coefficients, the table includes the (weighted) average coefficients (the reported standard errors account for both the variation between provinces, and their individual uncertainty). For both FE and PMG model we report standard errors in parentheses. For the Hausman test we report the χ^2-distributed test statistic (on top) and the p-value (in parenthesis). Correspondingly for the Fisher test (bottom two rows), we report the test statistic and the p-value in parentheses. ***, **, and * denote significance at the 0.01, 0.05 and 0.1 level, respectively

Table A.2 Hausman tests for the CPI effect

horizon	FE AR(1)	shock	PMG AR(1)	shock	Hausman
1	0.24976***	0.00030	0.25542***	0.00031	0.41
	(0.01796)	(0.00024)	(0.01548)	(0.00020)	(0.815)
2	0.08676***	−0.00017	0.09153***	−0.00016	0.16
	(0.01833)	(0.00024)	(0.01365)	(0.00015)	(0.923)
3	−0.11766***	0.00015	−0.12379***	0.00014	0.175
	(0.01832)	(0.00024)	(0.01090)	(0.00015)	(0.916)
4	−0.02760	−0.00252***	−0.03050**	−0.00251***	0.078
	(0.01774)	(0.00024)	(0.01279)	(0.00019)	(0.962)
5	−0.14035***	−0.00140***	−0.14539***	−0.00138***	0.338
	(0.01789)	(0.00024)	(0.01565)	(0.00014)	(0.844)
6	−0.10136***	−0.00102***	−0.10420***	−0.00101***	0.081
	(0.01806)	(0.00024)	(0.01342)	(0.00017)	(0.96)
7	−0.11175***	−0.00104***	−0.11444***	−0.00102***	0.057
	(0.01800)	(0.00024)	(0.01208)	(0.00016)	(0.972)
8	-3.9e−02**	-3.1e−05	-4.1e−02***	-3.7e−05	0.021
	(1.8e−02)	(2.4e−04)	(1.5e−02)	(2.0e−04)	(0.99)
Fisher-test					1.318
					(0.999)

Note: See Table A.1

Hausman test The Hausman test compares two estimators $\hat{\theta}_0$ and $\hat{\theta}_1$ for the same true parameter vector θ. The est is used to assess the validity of the assumption(s) underlying estimator $\hat{\theta}_1$.Under those assumptions both $\hat{\theta}_0$ and $\hat{\theta}_1$ are consistent, but only $\hat{\theta}_1$ is efficient. If the assumptions are violated $\hat{\theta}_1$ is inconsistent (rendering the comparison of efficiency irrelevant).

The test is most known for the comparison of fixed effects and random effects panel models (where random effects are efficient if the unit specific effects are uncorrelated to the regressors but inconsistent otherwise), or to test or endogeneity (where a potentially endogenous regressor renders the otherwise efficient OLS estimator inconsistent).

In our case, comparing the FE and the PMG estimator, the FE estimator is efficient if there is no slope heterogeneity since it retains more degrees of freedom, whereas the PMG estimator (allowing heterogeneity) is inefficient in the absence of this heterogeneity but always consistent.

Assuming the coefficient vectors are written as column vectors, the test statistic of the Hausman test is given by:

$$H = (\hat{\theta}_1 - \hat{\theta}_0)'(\hat{\Omega}_0 - \hat{\Omega}_1)^{+}(\hat{\theta}_1 - \hat{\theta}_0) \quad (6)$$

where $\hat{\Omega}_i$ is the variance-covariance matrix of $\hat{\theta}_i$, and the + superscript denotes a generalized inverse (since unlike the covariance matrices, the difference between covariance matrices is not necessarily invertible).

Table A.3 Hausman tests for the loan effect

horizon	FE		PMG		Hausman
	AR(1)	shock	AR(1)	shock	
1	0.39923***	−0.00713***	0.38375***	−0.00735***	0.073
	(0.01931)	(0.00066)	(0.04187)	(0.00056)	(0.964)
2	0.36785***	−0.00463***	0.33434***	−0.00496***	0.369
	(0.02004)	(0.00068)	(0.04114)	(0.00058)	(0.831)
3	0.30674***	−0.00377***	0.29142***	−0.00399***	0.297
	(0.02065)	(0.00070)	(0.03982)	(0.00057)	(0.862)
4	0.26546***	−0.00359***	0.24195***	−0.00380***	0.602
	(0.02070)	(0.00070)	(0.03960)	(0.00065)	(0.74)
5	0.25963***	−0.00384***	0.23092***	−0.00396***	0.123
	(0.02041)	(0.00069)	(0.03987)	(0.00063)	(0.94)
6	0.22759***	−0.00340***	0.20141***	−0.00347***	0.204
	(0.02078)	(0.00070)	(0.03712)	(0.00071)	(0.903)
7	0.20818***	−0.00654***	0.17572***	−0.00666***	0.883
	(0.02067)	(0.00069)	(0.03446)	(0.00071)	(0.643)
8	0.18195***	−0.00491***	0.15699***	−0.00500***	0.767
	(0.02117)	(0.00071)	(0.03261)	(0.00078)	(0.681)
Fisher-test					3.318
					(0.999)

Note: See Table A.1

Under the Null-hypothesis, H is χ^2-distributed with the degrees of freedoms matching the rank of $(\hat{\Omega}_0 - \hat{\Omega}_1)$, i.e., typically the number of coefficients that are compared.
Fisher-aggregation The Hausman test merely compares two models. However, in a local projections approach impulse responses are not generated by a single model but by separate models for every forecast horizon. In our case that means we have eight Hausman-tests when comparing a single set of impulse responses. The Fisher-method allows to test the joint Null-hypothesis that all individual Null-hypothesis are true for an arbitrary set of k tests by combining the p-values of the individual tests into a new test statistic:

$$S = -2 \sum_{i=1}^{k} \ln p_i, \tag{7}$$

which is χ^2-distributed with $2k$ degrees of freedom for independent tests. The test tends to overreject for dependent tests (as in our case), meaning we lean toward heterogeneity over homogeneity. In other words, we only enforce homogeneity if there is absolutely no evidence for meaningful heterogeneity at all.

Tables A.1–A.4 present the FE and PMG estimates for the four main variables, together with the corresponding Hausman test results.

Table A.4 Hausman tests for the property price effect

horizon	FE		PMG		Hausman
	AR(1)	shock	AR(1)	shock	
1	−0.0019	−0.0141***	−0.0059	−0.0117**	0.293
	(0.0180)	(0.0037)	(0.0239)	(0.0058)	(0.864)
2	0.0073	−0.0069*	0.0853***	−0.0080	12.777***
	(0.0177)	(0.0037)	(0.0281)	(0.0053)	(0.002)
3	0.0011	−0.0033	−0.0752**	−0.0038	11.169***
	(0.0175)	(0.0036)	(0.0295)	(0.0050)	(0.004)
4	−0.0091	−0.0072**	0.0409	−0.0062	12.933***
	(0.0168)	(0.0035)	(0.0262)	(0.0046)	(0.002)
5	0.0263*	−0.0050	0.0278	−0.0056	0.086
	(0.0157)	(0.0032)	(0.0260)	(0.0039)	(0.958)
6	−0.00619	−0.00025	0.00895	−0.00040	0.499
	(0.01558)	(0.00320)	(0.03147)	(0.00349)	(0.779)
7	−0.0155	−0.0147***	−0.0234	−0.0147***	0.429
	(0.0152)	(0.0031)	(0.0210)	(0.0040)	(0.807)
8	0.0311**	0.0077**	0.0025	0.0080*	208.106***
	(0.0147)	(0.0030)	(0.0160)	(0.0047)	(0)
Fisher-test					246.29***
					(0.000)

Note: See Table A.1

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