

# Interbank lending rates and monetary policy in China\*

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## Abstract

This paper explores the relationship between the Interbank lending rate CHIBOR and monetary policy in China. In addition to standard macroeconomic variables, a second model controls for the influence of stock prices. The empirical evidence suggests - based on monthly data in between April 1999 and November 2013 - that positive shocks to the Interbank lending rate lead to an increase of domestic credit in the short-run. Shocks to domestic credit have a persistent negative impact on the Interbank lending rate in China. Controlling for stock prices in a second model, yields comparable results for shocks to both CHIBOR and domestic credit.

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# 1 Introduction to the general idea

The Interbank lending market in the People's Republic of China has been subject to increasing fluctuations throughout the most recent years. In particular, potential liquidity problems of regional banks which have expanded their loans in the past have been the cause for rising concerns about the stability of the Chinese financial system.<sup>1</sup> Naturally, the Interbank lending rates, for example characterized by the 7-day Interbank repo rates, by the Chinese Interbank Offered Rate (CHIBOR) or by the Shanghai Interbank Offered Rate (SHIBOR), react to liquidity problems of banks. On June 5, 2013 Everbright & Co defaulted on interbank loans worth 6.5 billion Yuan - about \$ 1.07 billion - and in the subsequent weeks the Interbank lending rate as measured by the overnight CHIBOR rose sharply peaking at 15.3% on June 21. After the People's Bank of China (PBOC) decision to intervene on the money market, the overnight rate dropped to 6.5%. More recently - on December 20, 2013 - the Interbank lending rate in China rose again sharply and the central bank stepped in by injecting 300 billion yuan, around \$49.4 billion.<sup>2</sup>

Links between interbank lending rates, monetary policy and stock prices are not only important in China but are also observable in other countries around the world. The Lehman default in 2008 unarguably lead to shock waves in the financial market and Lehman's decision to file for chapter 11 bankruptcy protection on September 15, 2008 also lead to an increase in the Interbank lending rate as measured for the US by the effective federal funds rate. In the following months not only Fannie Mae and Freddie Mac were acquired by the federal government but also the FED took actions on the money market to provide more liquidity in the market followed by quantitative easing programmes. Similar movements could be observed by other central banks such as the European Central Bank or the Bank of England. While obviously the financial crisis must be interpreted multidimensionally and is impossible to relate to a single event, this paper takes these heavy daily fluctuations in the Interbank lending rate as motivation to explore the impact of the Interbank lending rate on monetary policy.

As a matter of fact, both the bankruptcy of Everbright & Co and the monetary actions taken by central banks during the financial crisis highlight the interdependence of Interbank

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<sup>1</sup>Descriptive information in this section are based on market information extracted from Bloomberg.

<sup>2</sup>Some analysts such as Min Shuai from Guotai Junan Securities Co pointed out that rising Interbank lending rates are due to the fact that banks need to meet reserve requirements with the PBOC. A future study might explore this with GARCH models.

lending rates and monetary policy decisions. Based on Vector autoregressions (VAR,) one might expect that domestic credit increases as response to positive shocks to Interbank lending rates.

In comparison with previous research, I will include a variable capturing Interbank lending rates in addition to standard macroeconomic variables which measure production, the exchange rate, inflation and monetary policy. The rationale is that shocks affecting the Interbank lending rate might have an impact on monetary policy decisions due to the mechanism of central bank interventions observed in event-studies. In fact, bank liquidity problems might lead to rising Interbank lending rates and require the central bank to step in the market. On the other hand, monetary decisions of the central bank clearly impact the Interbank lending rate from a theoretical point of view. This question will be addressed with VAR models placing the Interbank lending rate recursively before the monetary policy variable. Similar to the analysis of the interaction between monetary policy and stock prices conducted by (Bjørnland and Leitemo 2009), I would like to develop a specification capturing the interdependence between monetary policy and Interbank lending rates.

However, looking at financial data from Fall 2008, one can also conclude that stock price indices as measured for example by the S&P 500 decreased heavily. This provides motivation for including stock prices in the model. Previous research on monetary policy shocks included stock prices in the analysis and found significant interactions between both measures. A second model will be incorporated in the analysis for comparison purposes with previous research.

It is important to point out that considering monthly data will likely not reflect the huge short-term variations peaking for example in June 2013. The description of these events were solely meant to describe the motivation for a closer look at the Chinese Interbank lending market. A particular treatment of this question would require the use of more sophisticated methods, potentially by using high-frequency trading data. Therefore the main goal of this paper is to conduct an empirical investigation of Interbank lending rates and domestic credit in China which also includes stock prices in the analysis for comparison purposes.

The following paper is structured such that the subsequent section gives an overview over previous related research on monetary policy, then the specifics of the Chinese market (including a description of monetary policy in China and an overview over specific restrictions on the Chinese stock market) are pointed out. The next section is split up in a part about

the data sources and required steps of data preparation as well as a description of the general methods that will be applied for the two model specifications (including the motivation for each of them). Finally, the empirical results for the two model specifications will be presented and compared with previous research.

## 2 Literature review

As the paper alludes to many different aspects it is important to draw a clear distinction between the models in this paper and previous research to accentuate my proposed contribution. Previous research by (Bjørnland and Leitemo 2009) has concluded that there is a strong interdependence between US monetary policy and stock market returns. The empirical results suggest a real stock price decrease of 7 - 9 % due to a monetary policy shock of a 100 basis point increase of the federal funds rate. Vice versa, a real stock price increase of 1 % accounts for an interest rate increase of about 4 basis points. Related to this (Bernanke and Kuttner 2005) also underline stock market reactions to monetary policy shocks focusing on measuring unanticipated ('surprise') monetary contractions by the FED in an extension of the (Campbell and Ammer 1993) procedure. In particular, they explore the general impact of a 25 basispoint change in the interest rate on US stock markets. However, the paper mentioned before deal with the US market and rely on the federal funds rate as measure of monetary policy. In fact, this paper distinguishes itself from related research in two key characteristics: the focus on China and the approach to measure the interaction between Interbank lending rates with both monetary and stock price shocks.

Moreover, it is well-known that changes in monetary policy have an immediate impact on exchange rates. Expanding on the closed economy VAR literature such as (Sims 1980), (Bjørnland 2008) illustrates for the case of Norway that after entering a managed floating exchange rate regime both exchange rates and changes in asset prices measured by changes in stock prices on the Oslo Stock Exchange (OSEBEX) respond immediately to the release of surprising interest rate decisions. Opposed to this paper, Bjornland's model is based on a small open economy with floating exchange rates and is focused on the interaction between monetary policy and exchange rates.

Because of these research results exchange rates will be included as policy variables in the

analysis although for a subset of the time considered the exchange rate of the Renminbi (RMB) was fixed to the US-dollar. Furthermore, previous research by (Rigobon and Sack 2004) focused their analysis on the short-run effects of interest rates on the S&P 500 as well as on NASDAQ, and dealt with the effect of monetary policy on stock prices. However, the analysis was based on daily observations whereas this paper - due to the lack of daily data for China - is based on monthly data. Related research by (Lee 1992) or (Thorbecke 1997) also applied VAR models and lead to the conclusion that there is little interaction between monetary policy and stock prices while (Bjørnland and Leitemo 2009) found a stronger interaction between both terms. However, these paper focused on the interaction between stock prices and monetary policy solely. Therefore, I will also include a second model in this paper featuring an incorporation of stock prices as additional variable in the first model specification to measure the interdependence of monetary policy with both stock prices and Interbank lending rates. For this model, I will also again calculate the impulse responses to respective shocks. As a consequence I will also provide a comparison between the results from both models to gain insight into the interaction between Interbank lending rates, stock prices and monetary policy.

Furthermore, this paper does not relate to politically polarized discussions regarding the Renminbi (RMB) being undervalued and also does not aim at explaining the origin of China's current account surplus as in (Hoffmann 2013).

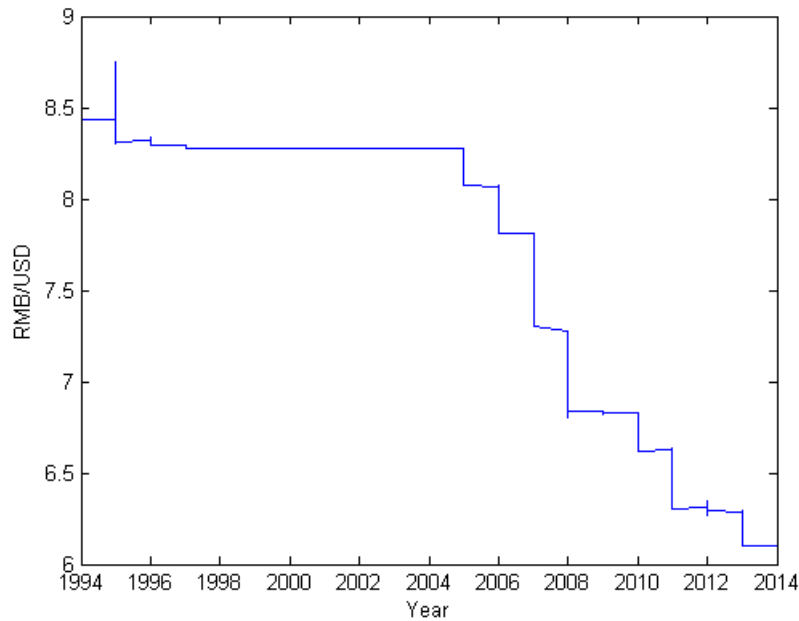
While in previous literature the emphasis lied on the interaction between monetary policy as measured for the US by the federal funds rate and stock prices, this paper will introduce Interbank lending rates as an additional variable and focus on China.

### **3 Monetary policy and the capital market in China**

As described in (McKinnon and Schnabl 2009), the relationship between the RMB/USD exchange rate can be divided into three different phases. Strictly speaking, before 1994 both currencies were unconvertible. Imports and exports were monopolized by state trading companies and official exchange rates were set. The second period covers the period from 1994 to July 2005 during which the exchange rate was fixed. In July 2005, the RMB was officially depegged from the USD although throughout the financial crisis restrictions were imposed and generally only fluctuations within a specific range are allowed. On figure 1 one can observe that from

1994 - 2014 the RMB appreciated from about 8.7 RMB/USD to 6.1 RMB/USD.

Figure 1: Exchange rate fluctuations RMB/USD from 1994 - 2014.<sup>3</sup>



Opening the Shanghai stock exchange in 1990 and the Shenzhen stock exchange in 1992, China slowly opened up to capital markets. However, the market is characterized by tight capital restrictions for foreign investors which are also reflected in a low Chinn-Ito-Index value of -1.1688 in 2011.<sup>4</sup> Ranking 110th out of 182 countries in the 2011 country ranking, one can clearly conclude that China is a country with low capital mobility. Individual foreign investors are for example restricted from buying A-shares on the Shanghai or Shenzhen stock exchange.<sup>5</sup> However, increasing the investment allowance of participants in the Qualified Foreign Institutional Investor (QFII) program that was launched in 2002 from US \$ 30 billion to US \$ 80 billion in April 2012 there are steps towards a better accessibility of the A-share market for foreign investors. These steps were continued in March 2014 by rising the shareholding limit for QFII from 20 % to 30 % and by doubling the daily trading band of the RMB.<sup>6</sup>

On November 9, 2008 - during the financial crisis - China passed a massive fiscal stimulus program of RMB 4 trillion - roughly \$ 586 billion. Moreover as described in (Han 2012) this

<sup>4</sup>Chinn-Ito-Index data is available for China in the period of 1984 - 2011. The index value in both 1984 and 2011 is -1.168826 and shows a low variation throughout the period with a lowest value of -1.8639 from 1987 - 1992.

<sup>5</sup>Only "Qualified Foreign Institutional Investors" are allowed to buy A-shares which are shares from in mainland-China based companies and are denominated in RMB.

<sup>6</sup>See (Takada 2014) for details on these changes. Available at <http://www.reuters.com/article/2014/03/20/us-china-foreigninvestment-stocks-idUSBREA2J02F20140320> (accessed on March 20, 2014).

fiscal stimulus program was also accompanied with monetary policy actions. For example, the loan quota mechanism which originally aimed at slowing down credit growth, were removed in 2007. Generally, the PBOC increased lending, and money supply as measured by M2 increased heavily.

Furthermore, it is a well-known fact that China has accumulated about \$ 3.8 trillion worth of international reserves. The significance of this amount and the opportunities for monetary policy interventions following from this - for example the option to support a particular exchange rate - give reason for including international reserves additionally to the exchange rate as variable in the model specifications. The prediction of the Mundell-Fleming model is hereby that in a country with low capital mobility and a fixed exchange rate international reserves decline in the short-run.

## 4 Theoretical model

Generally, China can be considered as a large open economy because it is integrated in world trade and a shock to the Chinese economy would likely have a significant impact on other economies who are for example dependent on exporting their goods to China (such as Germany). Moreover, China is a country with low capital mobility as can be seen from a low Chinn-Ito-Index. The exchange rate can be considered as a managed float. The Mundell-Fleming model as well the Dornbusch perfect-foresight extension (Dornbusch 1976) can be used to justify that because of heavy restrictions regarding the capital flow both the exchange rate and domestic credit can be treated as policy variables in the VAR specification (Mundell 1963). One of the (key) conclusions of the Mundell-Fleming model is the so-called "impossible trinity" which implies that a country cannot conduct an independent monetary policy, maintain a fixed exchange rate and allow capital to move freely at the same time (Obstfeld and Rogoff 1996). With China being a country of low capital mobility, i.e. without free capital movement, this provides theoretical grounding for treating the exchange rate and domestic credit as policy variables. The Mundell-Fleming will mainly be used as a point of comparison for reactions to changes in money supply. A stock market is not explicitly modelled in it.

An adequate representation could also imply a consideration of fixed exchange rates although recently the variation increases. Other approaches such as (Svensson 2000) that are

based on a new-Keynesian small open economy model which represents the Philipps curve do not seem to be adequate for China. Another option would be using a stylized New Open Economy Macroeconomics model similar to (Corsetti and Pesenti 2001) and (Obstfeld and Rogoff 2000).

The capital market restrictions imposed by China allow the PBOC to control both the money supply and the exchange rate. This provides reason to include both a measure of monetary policy and exchange rates in the VARs as policy variables. This can also be justified by the measures the PBOC took during the recent financial crisis, namely the changes regarding the loan quota mechanism mentioned in section 3. The general methodology resembles (Bjørnland and Leitemo 2009) but clearly distinguishes itself by the choice of the variables in each model specifications. In fact, the approach followed in this paper is to estimate two different models which solely distinguish by one variable on stock prices.

In the first approach I will measure monetary policy by domestic credit in log-first-difference form. Hereby, open market operations conducted by the PBOC for example in the aftermath of Everbright & Co's default in June 2013 can be measured better. Alternatively, one could include money supply based on the broad money definition M2. The latter could be justified by the fact that China implemented capital controls which allows them to control both the exchange rate and money supply. Building on this argument, one might argue that the broad definition of money supply as measured by M2 is a better measure for monetary policy. Hereby, M2 captures not only cash and checking deposits but also highly liquid assets such as money market mutual funds.

I will include data observations on both money supply M2 and domestic credit in the next section, conduct stationarity tests and will as robustness check in section 6.5 also conduct estimations with money supply. A final remark: the data on domestic credit has missing data observations in 1999 so that interpolation methods need to be applied again. Nevertheless, I will include domestic credit as measure for monetary policy in both models as it is better suited to capture open market operations conducted by central banks which - as the previous descriptions underline - play an important role in China. In section 6.5, I will replace domestic credit by money supply and therefore I also conduct stationarity tests on money supply in the subsequent section.

## 4.1 Data preparation

Table 1 summarizes potential variables, suggested data sources, the years of availability and the frequency of the data. Both SHIBOR, CHIBOR and the officially published Interbank lending rate from the PBOC could be used as proxy for Interbank lending rates. Moreover, both the Shanghai Stock Exchange Composite Index (SHCOMP) and the Shengzhen Stock Exchange Composite Index (SZCOMP) could be applied as proxy for stock market returns. Opposed to (Ho and Yeh 2010) who follow a similar set-up to measure the impulse responses to structural shocks in a small economy for the case of Taiwan I will use domestic credit as measure for monetary policy.<sup>7</sup> As the CHIBOR is available for a longer period and is a standard measure, I will include it in both models to capture the impact of Interbank lending rates.

Problematic is the fact that several data observations are missing. As a matter of fact

Variable	Source	Availability	Frequency
China Industrial production index	OECD	1999 - 2014	Monthly
CPI	OECD	Jan 1990 - Jan 2014	Monthly
Exchange rate RMB/USD	IFS	Feb 1984 - Jan 2014	Monthly or daily
Real effective exchange rate	BIS	Feb 1984 - Jan 2014	Monthly
International Reserves	Bloomberg or IFS	Oct 1995 - Jan 2014	Monthly
Interbank rates	PBOC	Jan 2002 - Dec 2010	Monthly
SHIBOR	Bloomberg	Jul 2009 - Jan 2014	Monthly or daily
CHIBOR	Bloomberg	Jul 1996 - Jan 2013	Monthly or daily
Stock Index: Shanghai SHCOMP	Bloomberg	Dec 1990 - Jan 2014	Monthly or daily
Stock Index: Shenzhen SZCOMP	Bloomberg	Jan 1992 - Jan 2014	Monthly or daily
Money Supply (MO, M1, M2)	Fred	Jan 1999 - Nov 2013	Monthly
Domestic credit	Bloomberg	Dec 1998 - Nov 2013	Monthly

Table 1: Overview over available variables and data sources.

given the data availability, it seems only feasible to estimate the VARs from April 1999 - November 2013 (equivalent to 176 observations on a monthly basis) if linear interpolations between missing observations can be applied. Indeed, data observations from CHIBOR are missing on standard statistical databases for July 1999 and February 2000. Moreover, the data on the industrial production index from OECD is missing in each January from 2006 - 2013: hence 8 observations are missing for this variable. The reason for this phenomenon is the fact that the China National Bureau of Statistics (NBS) has released a statement that reporting for

<sup>7</sup>Data on the Chinese money supply from September and November 2001 are missing. Potentially, because of monetary interventions in the aftermath of 09/11. Procedures on how to deal with missing data are summarized in (Howell 2007).

January is exempted officially. Because of this fact other proxies such as the China Value Added of Industry YoY (CHVAIOY) calculated by Bloomberg would not help in increasing the number of observations as they also officially do not report January observations for this indicator.<sup>8</sup> In addition to this, the China Industrial Production Index is considered as a standard proxy for GDP growth opposed to the CHVAIOY. I will also test for unit roots regarding money supply as I will replace domestic credit by money supply at a later point. One option to deal with missing data observations would be to run a simple OLS regression on past January observations and use the obtained estimates to predict the missing future January observations. However, due to a lack of previous observations this regression-based approach is not feasible for the Industrial production index from the OECD. Therefore, I conducted a simple linear interpolation between two dates for the missing January observations. For example the January 2006 observation is calculated in the following manner:

$$CP_{Jan,06} = \frac{CP_{Feb,06} - CP_{Dec,05}}{2}$$

where the subscripts stand for the month and the year of the respective observation. The absolute values obtained in this way will then be log-differenced for the main model specification.

## 4.2 Descriptive data

Table 2 summarizes descriptive data from April 1999 until November 2013 on Industrial production index, CHIBOR rate, SHCOMP index, exchange rate, the effective real exchange rate index calculated by the Bank for International Settlements (BIS), international reserves, money supply and domestic credit. Hereby, the mean, standard deviation, minimum and maximum are listed. In particular, the standard deviation might play an important role at the later part of the paper when impulse responses are calculated. One approach might consist of looking at impulse responses to a one or two standard deviation shock.<sup>9</sup> Remarkable is the large standard deviation from stock prices.

As the data on exchange rates is based on distinct exchange rate regimes I will also conduct

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<sup>8</sup>An official statement for this can be obtained for example from the description of the CHVAIOY which is available under: <http://www.bloomberg.com/quote/CHVAIOY:IND> (Accessed on February 22, 2014).

<sup>9</sup>The descriptive data is obtained from the MATLAB file *Descriptive\_data.m* which uses the function *Sum\_stats.m*.

Variable	Minimum	Maximum	Mean	Standard deviation
China Industrial production index	102.3	123.2	113.38	3.8109
Exchange rate	6.0918	8.2828	7.5155	0.8183
Real effective exchange rate index	84	113.76	96.193	7.1511
International Reserves	1.4667 e <sup>5</sup>	3.7895 e <sup>6</sup>	1.4155 e <sup>6</sup>	1.2041 e <sup>6</sup>
CHIBOR	0.72	6	2.375	0.95148
Stock Index: Shanghai SHCOMP	254.47	1532.7	710.39	335.13
Money supply: M2	1.0915 e <sup>13</sup>	1.0893 e <sup>14</sup>	4.2828 e <sup>13</sup>	2.9041 e <sup>13</sup>
Domestic credit	9865.2	90327	35561	23133.75

Table 2: Descriptive data: absolute data

a test for structural breaks associated with these regime changes. (Thoma 2007) provides a discussion of the impact of structural changes on the VAR lag length as measured by changes in the parameter on the output gap term.

Possible tests for structural breaks include the Chow test (Chow 1960) where one might suspect one structural break for the exchange rate. In fact, I am not able to reject the null hypothesis that coefficient estimates for subsamples of the time series are the same. This means that the Chow-test suggests that there is no structural break in the exchange rate. One problem that arises is that the change in exchange rates is likely associated with an increase in the variance of changes in the exchange rate. Despite the result of the Chow-test I therefore test for changing variances. On this purpose I split up the data observations in two samples: the first from April 1999 until the change in exchange rate regime in July 2005 and a second from August 2005 until November 2013.

Time	Variance
April 1999 - November 2013	0.6696
April 1999 - July 2005	0.0004
August 2005 - November 2013	0.4031

Table 3: Variance of subsamples

Not surprisingly there is a huge change in the variance of the exchange rate through the change in the exchange rate as can be seen from table 3. For that reason I decide to include the effective real exchange rate index as measured by the Bank for International Settlements (BIS) which is not affected by the change in exchange rate regime.

### 4.3 Stationarity tests

In the first step, it must be determined if the underlying time series' are stationary to be able to apply VAR. I will apply the standard Augmented Dickey-Fuller-test (ADF) to determine stationarity.

As a matter of fact evaluating figure 2 showing the changes in CHIBOR rate and figure 3 representing the stock index SHCOMP, one can get the impression that the underlying time series is not stationary. Conducting the augmented Dickey-Fuller (ADF) test (Dickey and Fuller 1979) on absolute variable values lasting from April 1999 until November 2013 I am unable to reject the null hypothesis of a unit root for the Industrial production index, for international reserves, CHIBOR, SHCOMP, domestic credit and M2.<sup>10</sup> The failure to reject the null hypothesis of a unit root is a typical behavior observed in financial time series of prices.

Figure 2: CHIBOR rates from 1996 - 2014

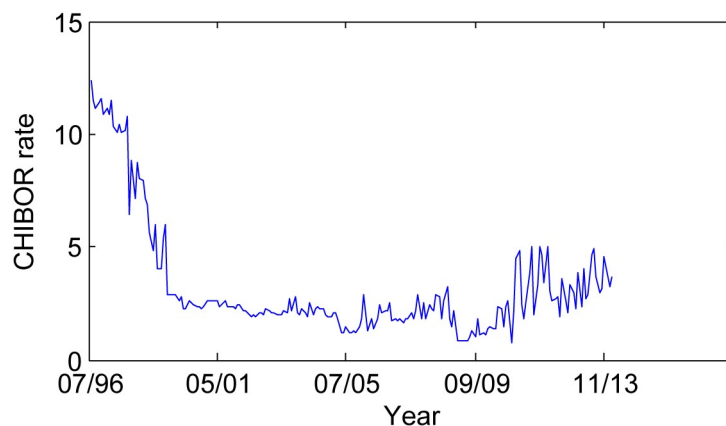
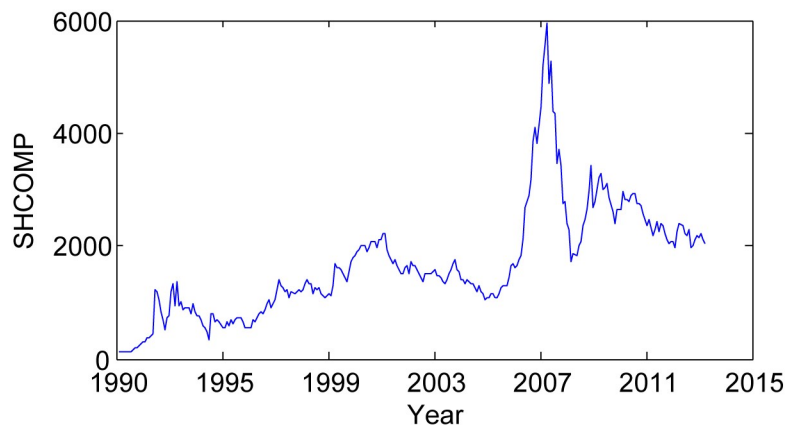


Figure 3: SHCOMP from 1990 - 2014



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<sup>10</sup>The subsample was chosen such due to the problems with missing data mentioned before. Simple linear interpolations are conducted for two missing observations of the CHIBOR variable as well as for 8 missing observations of the Industrial production index.

It is a prerequisite for VARs that the variables are stationary following for example (Bjørnland and Leitemo 2009). Considering logs of all absolute values and conducting ADF-tests for a log-specification, one fails to reject the null hypothesis of a unit root on all variables.<sup>11</sup> In fact, the null can only be rejected for the real effective exchange rate index and CHIBOR only. This means that for example the unit root null hypothesis cannot be rejected for the log of the industrial production index which differs from (Bjørnland and Leitemo 2009). In a next step I also considered log-differenced variables where I end up rejecting the null hypothesis of a unit root - again based on an ADF test. Hereby, the effective real exchange rate is included because of the the ambiguous results of the Chow-test.

An overview of the final model results of the augmented Dickey-Fuller test is also summarized in table 4.

Variable	Null hypothesis	P-value	Statistic
China Industrial production index	Rejected	0.0010	-21.6913
CPI	Rejected	0.0010	-9.2462
Exchange rate	Failure to reject	0.9990	14.3528
Real effective exchange rate	Rejected	0.0010	-9.0959
International Reserves	Rejected	0.0010	-8.8951
CHIBOR	Rejected	0.0010	-18.1680
Stock Index: Shanghai SHCOMP	Rejected	0.0010	-12.1787
Money supply: M2	Rejected	0.0010	-5.2181
Domestic credit	Rejected	0.0010	-7.5419

Table 4: Results from ADF-test on log-differenced variables with a critical value of -1.9424 for the ADF-statistic (Cheung and Lai 1995). Significance level of  $\alpha = 0.05$ .

## 4.4 Final model specification

Based on the results of the stationarity tests there exist two possibilities of the model specification: on the one hand a log-first-difference form or on the other hand a form based on relative changes on each variable. In this paper I follow the former option where I include the effective real exchange rate. The prerequisite of stationarity is fulfilled for both potential versions for all variables.

After selecting the measurement type of the input variables table 5 summarizes the input parameter for each model specification as well as the measurement for each of these variables.

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<sup>11</sup>Empirical results and the code are documented in the Appendix or available upon request.

Money supply is not included in this summary as it will only be included in the robustness section. Recall that my goal consists of estimating two models that solely differentiate in the stock index variable.

Variable	Model 1	Model 2	Notation	Measurement
China Industrial production index	X	X	IP	Log-first-difference
CPI	X	X	$\pi$	Log-first-difference
Effective real exchange rate	X	X	Ex	Log-first-difference
International Reserves	X	X	IR	Log-first-difference
CHIBOR	X	X	CH	Log-first-difference
Stock Index: Shanghai SHCOMP		X	SH	Log-first-difference
Domestic credit	X	X	DC	Log-first-difference

Table 5: Variable input for both models.

In both models, inflation as measured by the Consumer Price Index (CPI), a proxy for GDP growth, the effective real exchange rates, international reserves, data on the CHIBOR and money supply will be included. Motivated by (Bjørnland and Leitemo 2009) as well as by (Bernanke and Kuttner 2005) I will include data on the SHCOMP as measure of stock price reactions in the second model. Additionally, China has accumulated foreign reserves of about \$ 3.8 trillion in 2013 and I will include them as variable. The intuition is that these can be used for market interventions to support the managed float of RMB. Conventional expectations for a country with low capital mobility are that an expansionary monetary policy leads to a decline in international reserves when the central bank uses the international reserves to maintain a fixed exchange rate regime (sterilization). The reason for estimating two models that only differ in one variable on stock prices is that this allows for comparison with results obtained in previous research on the interaction between stock prices and monetary policy.

## 5 VAR method

Introduced by (Sims 1980) Vector autoregressions (VAR) became a powerful tool in the analysis of the impact of economic policies on macroeconomic variables. In general there exist three different forms of VAR estimations: reduced form, recursive and structural form. As mentioned before the methodological approach in this paper will be based on the moving average version of the VAR where the current values of each variable is solely dependent on the

residuals. Generally, two steps need to be followed to estimate the recursive version of a VAR (Hamilton 1994):

1. The reduced forms must be estimated for all equations which form the model.
2. The Choleski-decomposition of the variance-covariance matrix of the residuals must be estimated.

Standard OLS cannot be applied to estimate the general VAR model as the independent variables will be correlated with the error term, violating the classical linear regression assumptions. As a consequence OLS estimates would be inconsistent. However, one can obtain the reduced form of the VAR model and follow the steps outlined above to estimate the structural parameters. With  $Z(t)$  containing the macroeconomic variables the moving average form of the general VAR model can be expressed in the following manner without deterministic terms:

$$Z_t = B(L)u_t$$

In this specification  $u_t$  captures the reduced-form residuals and  $B(t)$  is the  $(n \times n)$  matrix featuring the lag operators. Following (Bjørnland and Leitemo 2009) one can express the VAR then in the following way with  $\epsilon_t$  representing the structural disturbances:

$$Z_t = B(L)S\epsilon_t \tag{1}$$

As usual in recursive VAR representations, the ordering of the variables will be essential for the interpretation. Therefore, for each model I will provide an economic reason for the ordering chosen. The abstraction from deterministic terms in both models follows (Bjørnland and Leitemo 2009).

Following (Greene 2008) I will follow standard assumptions for VAR estimations such that I assume the  $x_t$  to be weakly time exogenous. This implies that an estimation independently of the marginal distribution of  $x_t$  is possible while at the same time Granger causality is not excluded between  $x_t$  and  $y_t$ . Moreover, I run Granger causality Wald tests for both models.<sup>12</sup>. The null hypothesis in these tests is that the variable  $x_t$  does not Granger-cause  $y_t$ .

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<sup>12</sup>The empirical output is available upon request.

In summary, the results suggest that individually stock prices Granger-cause Industrial production. Additionally, domestic credit, international reserves and the real effective exchange rate Granger-cause consumer prices. International reserves Granger-causes CHIBOR and CHIBOR Granger-causes domestic credit at a significance level of  $\alpha = 10\%$ . In model 1, the real effective exchange rate Granger-causes Industrial Production Index, international reserves Granger-causes consumer prices and CHIBOR, the real effective exchange rate and domestic credit Granger-cause international reserves and CHIBOR Granger-causes domestic credit at a significance level of  $\alpha = 10\%$ . However, with the standard assumption of weakly exogeneity as in (Greene 2008) the Granger-causal relationships do not impose a problem for the VAR estimation.

The effect of shocks to the system will be calculated such that innovations for specific variables will be introduced for one period and are set equal to zero in the subsequent periods. The impulse response will then show the path back to equilibrium and the impact on other variables.<sup>13</sup>

The number of lags for the VAR model is typically determined by using either likelihood ratio test (for example following the statistic proposed by Sims (1980)), the Akaike information criterion or the Schwarz-Bayesian criterion. Similar to standard literature I will use the Akaike information criteria and check for robustness with different lag numbers in section 6.5.

## 5.1 First model

The variables in the first model will likely be ordered such that the CHIBOR variable is placed second-last, right before domestic credit. In the first model specification the ordering of the structural shocks - putting domestic credit shocks last - looks in the following manner:

$$[\epsilon_t^{IP}, \epsilon_t^{\pi}, \epsilon_t^{CH}, \epsilon_t^{IR}, \epsilon_t^{Ex}, \epsilon_t^{DC}]'$$

The indices are denoted in the following way: 'IP' stands for industrial production, ' $\pi$ ' for inflation, 'CH' for CHIBOR, 'IR' for international reserves, 'Ex' for real effective exchange rates and 'DC' for domestic credit. The key interpretation will aim at the impulse responses to shocks to CHIBOR  $\epsilon_t^{CH}$  and to shocks to domestic credit  $\epsilon_t^{DC}$  as measure of monetary policy. The ordering resembles (Bjørnland and Leitemo 2009) and (Ho and Yeh 2010). As discussed

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<sup>13</sup>See (Greene 2008) pages 586-602 for details.

before the real effective exchange rate is considered as a policy variable because China is a country with low capital mobility. Hereby, international reserves are placed before the real effective exchange rate. The intuition for this is again that international reserves can be used by the PBOC to influence the exchange rate. Moreover, it follows standard procedures such that macroeconomic variables do not change simultaneously with policy variables. As a consequence, the first model specification based on the general form in (1) takes the following form with  $Z_t = [\Delta IP_t, \Delta CP_t, \Delta CH_t, \Delta IR_t, Ex_t, \Delta DC_t]'$ :

$$\begin{pmatrix} \Delta CP_t \\ \Delta \pi_t \\ \Delta CH_t \\ \Delta IR_t \\ Ex_t \\ \Delta DC_t \end{pmatrix} = B(L) \begin{pmatrix} S_{11} & 0 & 0 & 0 & 0 & 0 \\ S_{21} & S_{22} & 0 & 0 & 0 & 0 \\ S_{31} & S_{32} & S_{33} & 0 & 0 & 0 \\ S_{41} & S_{42} & S_{43} & S_{44} & 0 & 0 \\ S_{51} & S_{52} & S_{53} & S_{54} & S_{55} & 0 \\ S_{61} & S_{62} & S_{63} & S_{64} & S_{65} & S_{66} \end{pmatrix} \begin{pmatrix} \epsilon_t^{IP} \\ \epsilon_t^{\pi} \\ \epsilon_t^{CH} \\ \epsilon_t^{IR} \\ \epsilon_t^{EX} \\ \epsilon_t^{DC} \end{pmatrix}$$

In the next step reasonable restrictions must be determined for  $B(L)$  before the impulse responses can be calculated. I hereby follow a standard approach in the literature and assume zero short-run restrictions for the structural VAR estimation. Another approach would be to follow the Blanchard-Quah method (Blanchard and Quah 1989) which involves imposing zero long-run restrictions. In fact, with the Blanchard-Quah decomposition real and nominal shocks are identified but they do not have long-run effects on first differences of the variables. However, adhering to the zero short-run restriction assumption is closer to (Bjørnland and Leitemo 2009) and I will therefore follow this approach in this paper.

## 5.2 Second model

When imposing restrictions on the variables the ordering of the structural shocks in the second model is specified in the following way:  $[\epsilon_t^{IP}, \epsilon_t^{\pi}, \epsilon_t^{CH}, \epsilon_t^{SP}, \epsilon_t^{IR}, \epsilon_t^{EX}, \epsilon_t^{DC}]'$ .

where 'SP' stands for stock prices as measured by the SHCOMP stock index. In comparison with the first model, I added shocks to stock prices  $\epsilon_t^{SP}$ . Moreover, I assume that Interbank lending rates react faster to shocks than stock prices by placing this variable before stock prices in the VAR representation. This assumption might be questionable and relates to the question

if adjustments on Interbank lending market or on stock markets are faster. However, I also tested the model with placing SP before CHIBOR and the results do not change qualitatively. The second model based on the general form in (1) with  $Z_t = [\Delta IP_t, \Delta CP_t, \Delta CH_t, \Delta SP_t, \Delta IR_t, Ex_t, \Delta DC_t]'$  will then be:

$$\begin{pmatrix} \Delta CP_t \\ \Delta \pi_t \\ \Delta CH_t \\ \Delta SP_t \\ \Delta IR_t \\ \Delta Ex_t \\ \Delta DC_t \end{pmatrix} = B(L) \begin{pmatrix} S_{11} & 0 & 0 & 0 & 0 & 0 & 0 \\ S_{21} & S_{22} & 0 & 0 & 0 & 0 & 0 \\ S_{31} & S_{32} & S_{33} & 0 & 0 & 0 & 0 \\ S_{41} & S_{42} & S_{43} & S_{44} & 0 & 0 & 0 \\ S_{51} & S_{52} & S_{53} & S_{54} & S_{55} & 0 & 0 \\ S_{61} & S_{62} & S_{63} & S_{64} & S_{65} & S_{66} & 0 \\ S_{71} & S_{72} & S_{73} & S_{74} & S_{75} & S_{76} & S_{77} \end{pmatrix} \begin{pmatrix} \epsilon_t^{IP} \\ \epsilon_t^{\pi} \\ \epsilon_t^{CH} \\ \epsilon_t^{SP} \\ \epsilon_t^{IR} \\ \epsilon_t^{Ex} \\ \epsilon_t^{DC} \end{pmatrix}$$

Again the standard approach in literature will be followed with zero short-run restrictions.

## 6 Empirical results

This section is structured such that for respective shocks both models are compared directly. This involves a conclusion which model seems to fit better taking into account results from previous research. The exact MATLAB code is originally based on the VAR toolbox from Ambrogio Cesa-Bianchi (Cesa-Bianchi 2014) which also includes functions from the Econometrics toolbox by James P. LeSage (LeSage 2014). However, I modified functions contained in these toolboxes. Extensive output summaries as well as the modifications conducted are available upon request for the purpose of replicating my results.

The main results include a comparison of impulse responses of shocks to Industrial Production, CHIBOR, stock prices as measured by SHCOMP and domestic credit between both models and to previous research. One standard routines in the literature is to consider a shock of one standard deviation to each of the repective variable. Note that this implies for the interpretation that I am considering positive shocks. Because of illustrative reasons I will consider unit shocks instead. Results go through for considering shocks of one standard deviation.

Moreover, it is assumed that there are no short-run restrictions allowing for a standard Choleski decomposition. An alternative approach would be the to impose sign restrictions

on the impulse responses following (Uhlig 2005) or to use a modified algorithm proposed by (Fernández-Villaverde and Watson 2005).

Table 6 summarizes the output respectively for a maximum number of 10 lags. Based on the Akaike information criterion the lag number with the lowest AIC-value is selected. This means that three lags are selected for the first model and two for the second model.

Number of lags	Model 1	Model 2
1	-31.2608	-33.3303
2	-31.4941	-33.5408
3	-31.5546	-33.4917
4	-31.4198	-33.3719
5	-31.4112	-33.4260
6	-31.4598	-33.5034
7	-31.4615	-33.4797
8	-31.3477	-33.3436
9	-31.2057	-33.1536
10	-31.2545	-33.2277

Table 6: Variable input for both models.

In both models I check that the VAR is invertible and stable. The stability condition implies that the VAR process generates stationary time series and is fulfilled for both models. Under the standard VAR notation this means for a  $Var(p)$  model that:

$$\det(I_n - A_1z - A_2z^2 - \dots - A_pz^p) \neq 0 \text{ for } z \leq 1 \quad (2)$$

The invertibility of a VAR model in the moving average form is determined by the following condition whereby the notation introduced before is followed ((Lütkepohl 2005)):

$$\det(I_n + B_1z + A_2z^2 + \dots + B_qz^q) \neq 0 \text{ for } z \leq 1 \quad (3)$$

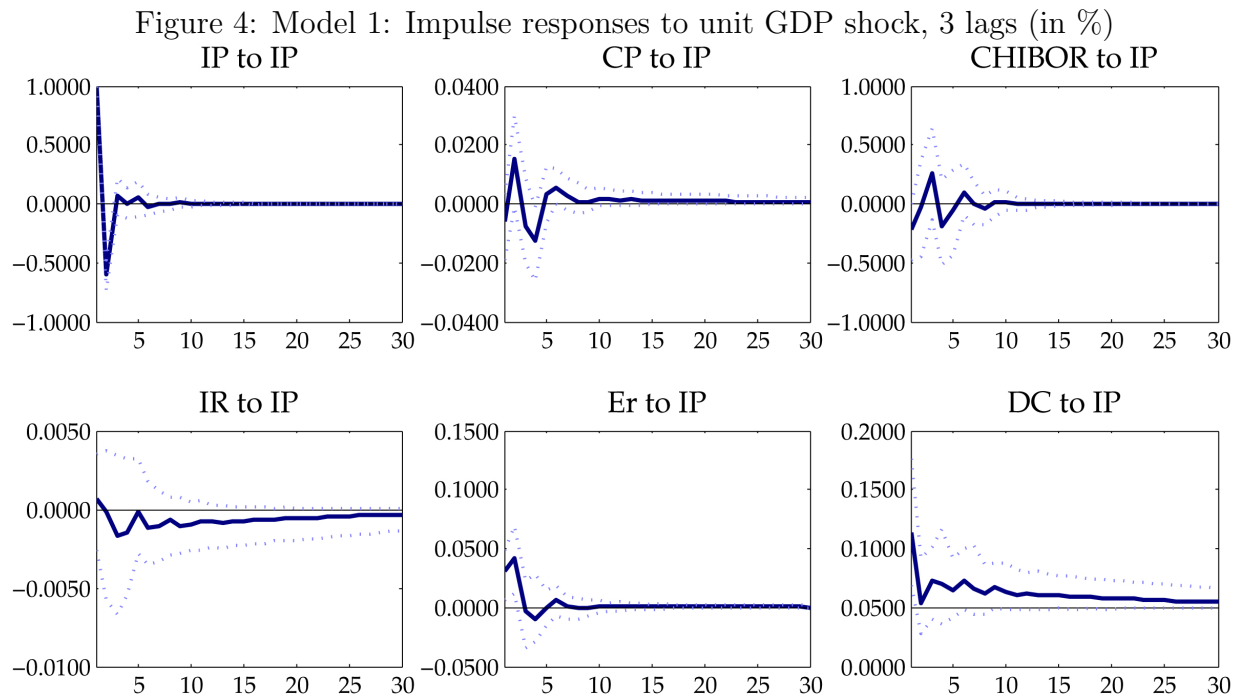
The graphs will depict the impulse responses up to 30 months after the shock occurs and show the percentage effect.

The results for this paper are based on a first difference log specification in which monetary policy is measured by domestic credit and the effective real exchange rate from the Bank

for International Settlement (BIS) is included.<sup>14</sup> For the purpose of completeness the graphs will include the responses of all variables - i.e. including the one which is ‘shocked’. Especially, the interpretation for the latter variable will be limited and must be done carefully. Once again including all impulse responses serves the purpose of complete reporting.

Moreover, after a comparison of the results obtained from each of the two models robustness checks are summarized in 6.5. Hereby, the lag length is alternated, in one version domestic credit is replaced by money supply (M2) and finally a specification based on relative changes of the underlying variables (i.e. without taking logs) is summarized. The latter specification would be another measurement option as alluded to in 4.3 based on stationarity tests. Details of the robustness tests are relegated to the Appendix.

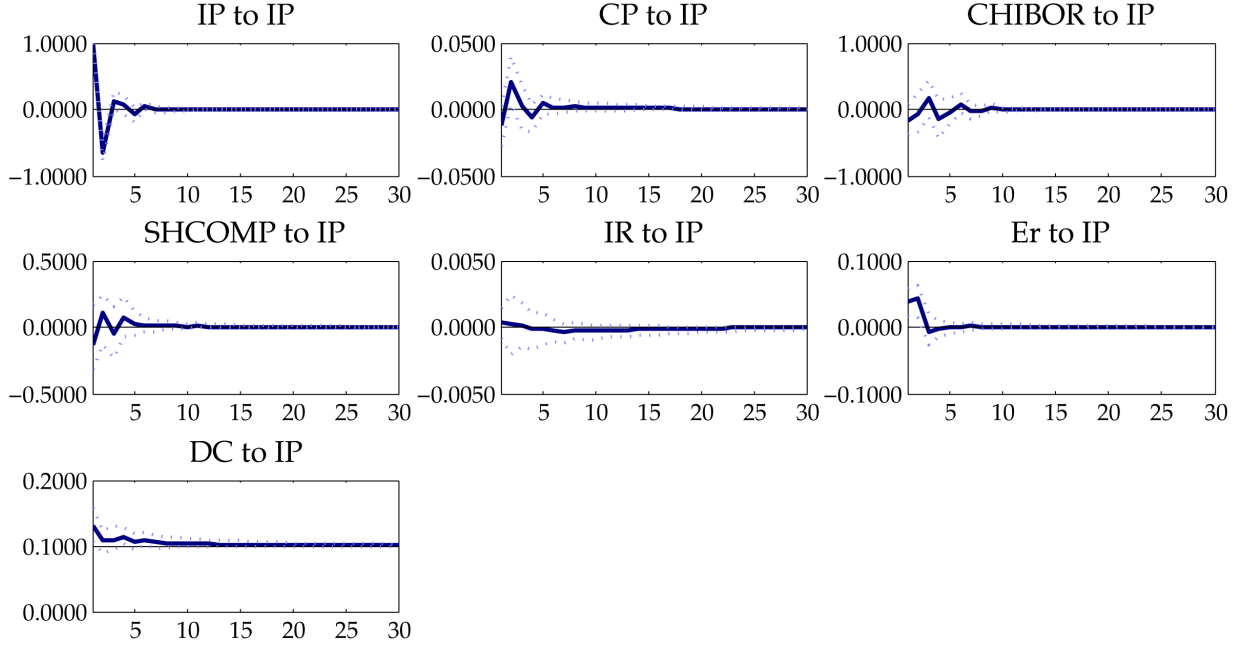
## 6.1 Shocks to Industrial Production



In both models impulse responses to unit shock of the Industrial Production Index react in a similar pattern for consumer prices, CHIBOR, real effective exchange rates and domestic credit. However, there are differences in the magnitude of the responses. While CHIBOR growth rate slows down by about 0.3 % as a response to a unit shock to Industrial Production index in model 1 the negative effect is less strong in model 2 directly after the shock. Until

<sup>14</sup>The stationarity for these variables was already verified in section 4.3.

Figure 5: Model 2 Impulse responses to unit GDP shock, 2 lags (in %)



five months after the shock the real effective exchange rate shows a positive response to shocks to Industrial production before the effect dies out. Moreover, there is a persistent effect on domestic credit growth of 0.05 % in model 1 and 0.1 % in model 2 which is significantly positive.

To put it in a nutshell, regarding the impulse responses to shocks to Industrial Production both models offer only slight differences. The effect on stock prices in model 2 is slightly negative in the first month after the shock but turns positive thereafter. Moreover, the effective real exchange rate appreciates in both models whereby the effect is slightly stronger in model 2.

## 6.2 Shocks to CHIBOR

It is interesting to see from figures 6 and 7 that reactions of domestic credit to CHIBOR rate shocks do not become less severe after incorporating stock prices in the model. Impulse responses from Industrial Production increase five months after the shock from -0.001 % five months after shock to about 0 % after including stock prices. A similar qualitative observation can be made regarding international reserves. Although basically these effects are diminishingly small this gives some support for the second model. Stock prices itself decrease which is an intuitive result taking Interbank lending rates as proxy for the overall interest rate level in an economy.

Figure 6: Model 1 Impulse responses to unit CHIBOR shock, 3 lags (in %)

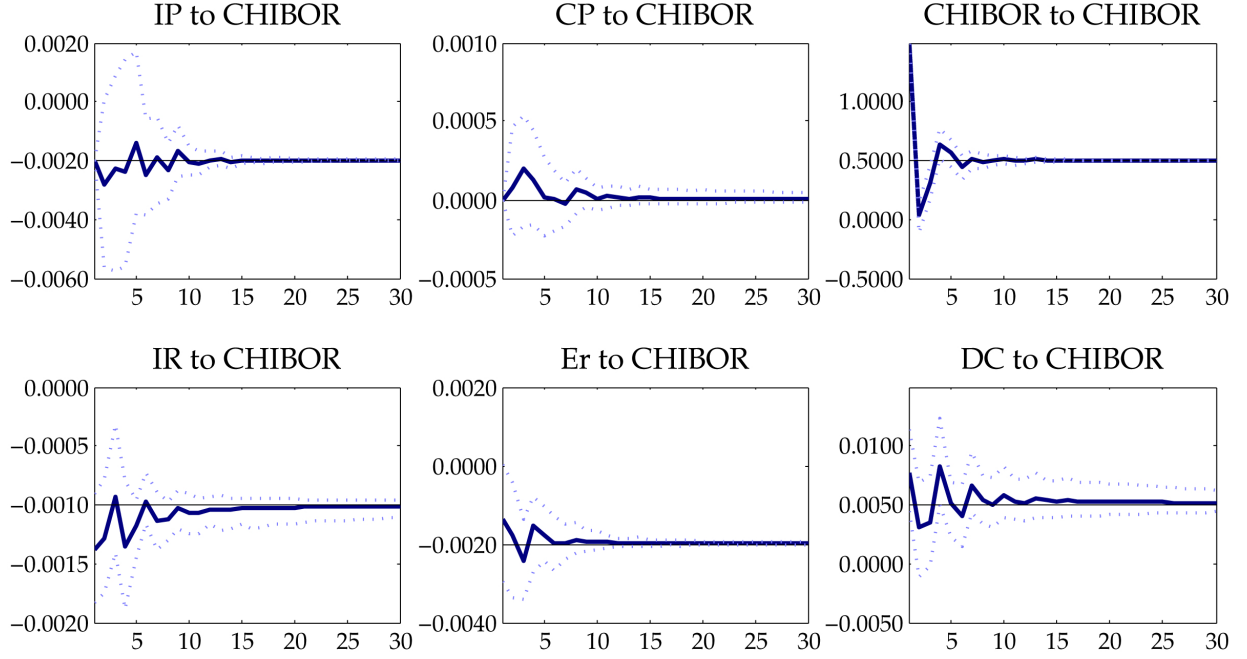
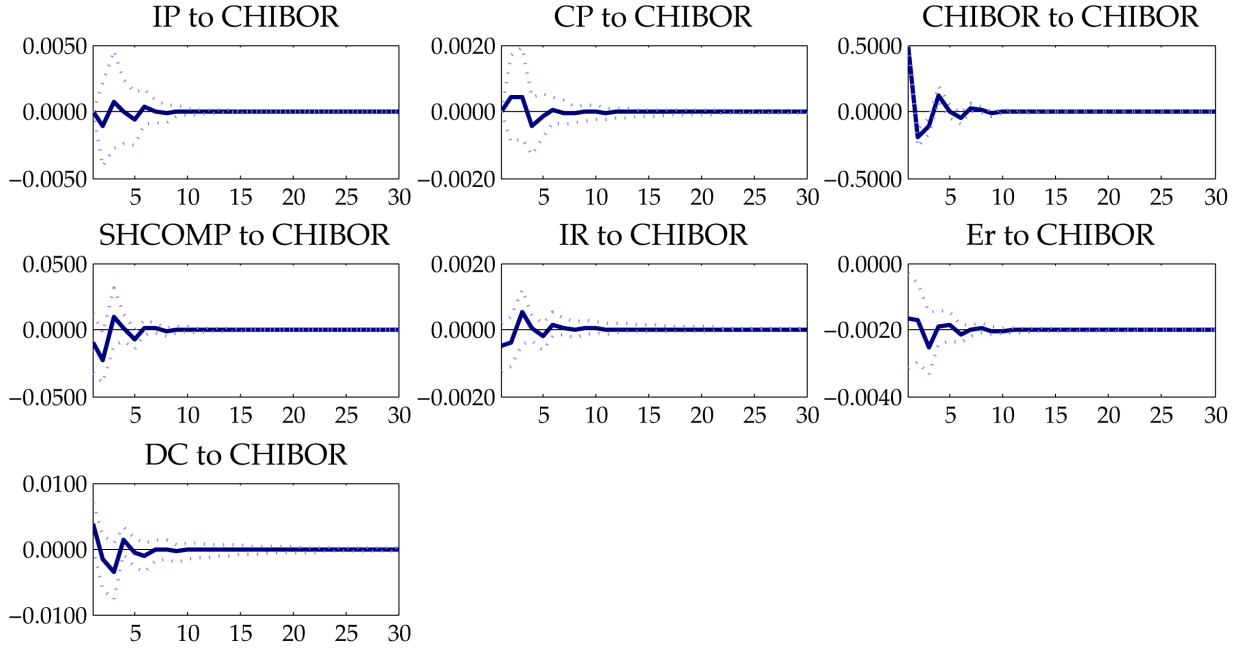


Figure 7: Model 2 Impulse responses to unit CHIBOR shock, 2 lags (in %)



The effective real exchange rate decreases in both models after a shock whereby this effect is persistent. The standard intuition would be that if interest rates increase, the exchange rate RMB/USD would appreciate and discords to it. Obviously, this relates to the fact that China had a fixed exchange rate regime for over half of the time considered. Domestic credit reacts in a slightly positive way to a positive CHIBOR shock. Basically, that's exactly the effect we were trying to explore empirically as an explanation for the reaction of the PBOC after the default of Everbright & Co (outlined in section 1). However, the effect bounces around, alternates in sign and is generally of a small magnitude. Therefore, one must be really careful before drawing the conclusion that spikes in Interbank lending rates lead to increases in domestic credit - particularly if thinking of the general increasing trend of domestic credit over time. In fact, the central bank could also follow a contractionary monetary policy as response to spikes in Interbank lending rates to further cope with inflation. This could play a role considering the fact that as a first response to a rising CHIBOR rate inflation also increases.<sup>15</sup> As a matter of fact, the positive impulse response of domestic credit is significant in the first model, while for the second model the sign is not statistically significant at the 5 % level.

Figure 7 confirms the doubts about the robustness of the results with respect to the impulse of domestic credit. In this figure one can see the upper and lower bands of the impulse which are based on bootstrapping with 100 draws. As a matter of fact the upper band is positive while the lower one is negative. Therefore, the significance of the results regarding the impulse response of domestic credit to CHIBOR shocks is ambiguous and economically in this sample close to zero. The theoretical expectation of an increase in domestic credit derived from an event-study is hence only partially reflected empirically. However, the positive response immediately after occurrence of a CHIBOR shock could be a sign that using models that are based on higher frequencies (for example: daily data observations) could address the question of an expansion in domestic credit as response to positive shocks to CHIBOR better.

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<sup>15</sup>Remark: one might argue that a Vector Error Correction Model (VECM) would be more appropriate if there are cointegrating relationships. The Johansen test (Johansen 1991) suggests that the time series are cointegrated of order 6 which means that one would need to take six differences to obtain a covariance stationary time series. However, my model specification is based on stationary log-differenced variables rather than non-stationary variables that with a cointegrating vector could form a stationary linear combination. Therefore I only report the outcome of the Johansen test at this point in time for completeness. In any case, such a VECM specification without trend leads to nearly the same results regarding the impulse responses of domestic credit to positive CHIBOR shocks as can be seen in figure 16. In this sense it does not provide better empirical evidence for the theoretical expectation of rising domestic credit as response to positive CHIBOR shocks. In conclusion, I include VAR estimations in the main part because in this way one can maintain more observations and because my variables are stationary.

### 6.3 Shocks to domestic credit

Figure 8: Model 1 Impulse responses to unit domestic credit, 3 lags (in %)

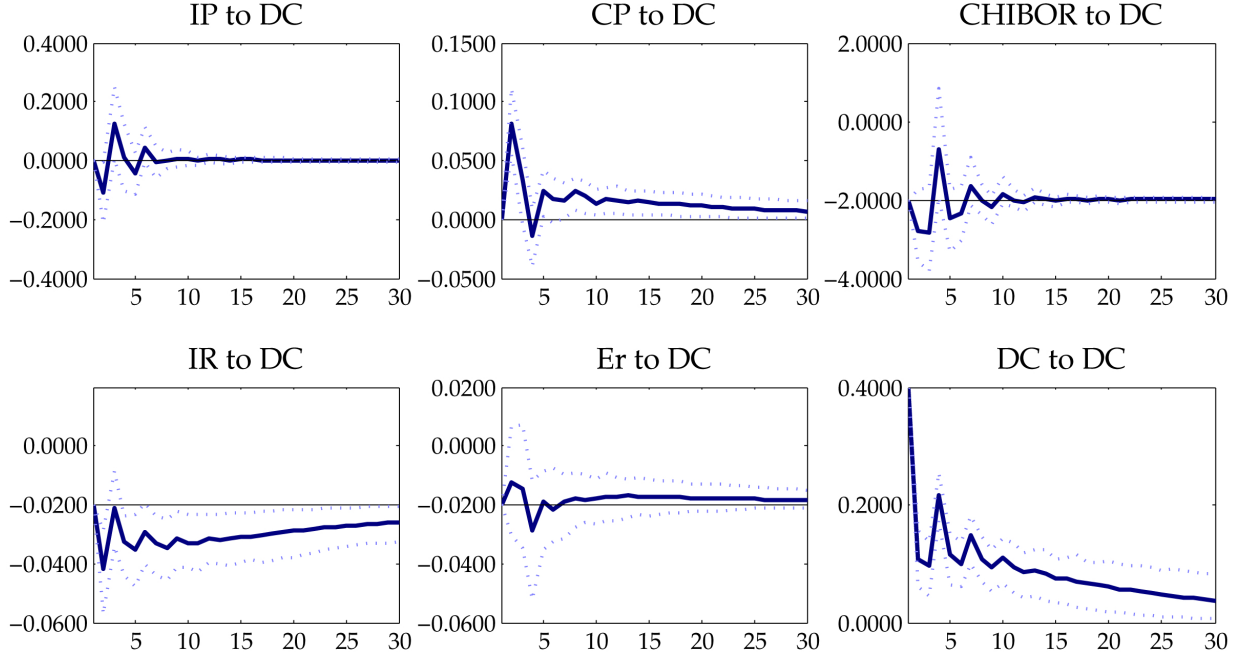
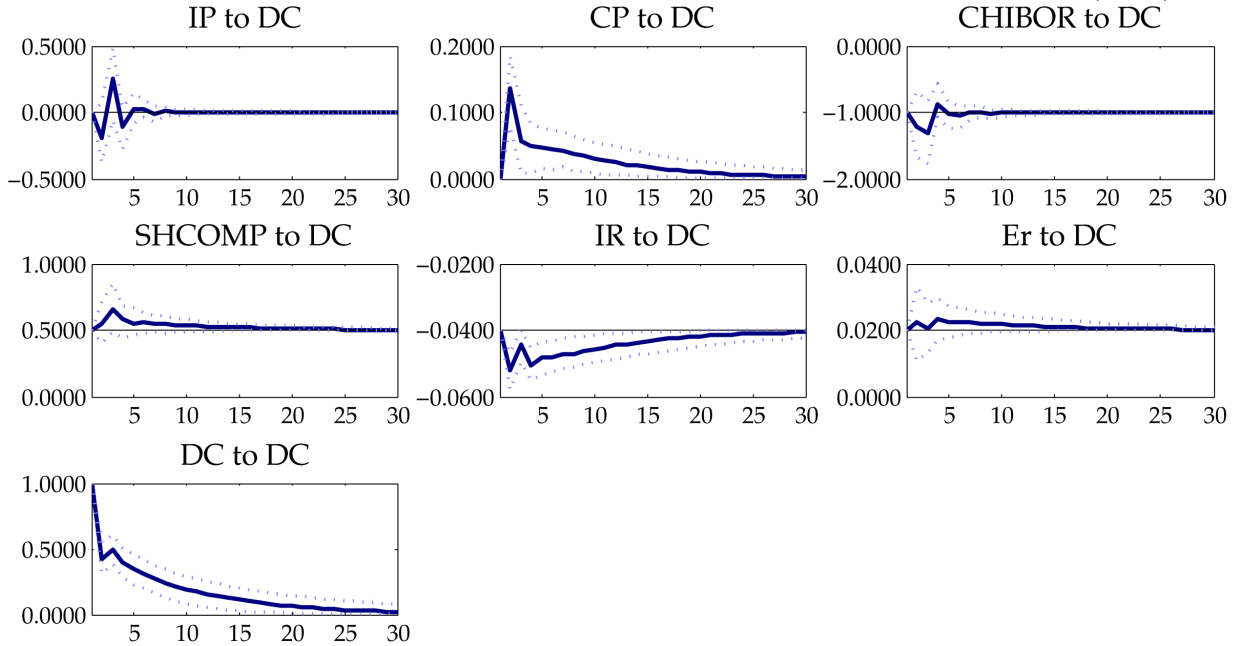


Figure 9: Model 2 Impulse responses to unit domestic credit shock, 2 lags (in %)



The impact of shocks to domestic credit on consumer prices, international reserves and the exchange rate is less severe in model 2. A (positive) shock to domestic credit translates into an increased growth of stock prices in figure 9 peaking at about 0.65 % 4 months after the shock which is statistically significant at the 5 % level. This is an intuitive result and distinguishes from (Bjørnland and Leitemo 2009) whose positive response of stock prices to a contractionary

monetary policy shock is counterintuitive.

At the same time one can interpret the graph at the bottom of figure 9 such that after a shock to money supply the central bank will inject more money into the market, though at a very small percentage level.

A general observation is that the CHIBOR falls persistently after a shock to domestic credit in both models. This is consistent with the effect of an increase in money supply in the Mundell-Fleming model. In model 2, one can see that the long-run effect of a domestic credit shock is  $-2\%$  while in model 1 it is only  $-1\%$ . The persistence is contradictory to the Mundell-Fleming model which predicts a gradual adjustment back to the previous interest rate level. It is remarkable that in both models there is no evidence for the so-called liquidity puzzle. The liquidity puzzle exists if monetary policy does not have a negative correlation with short-term nominal interest rates. However, taking CHIBOR as measure for short-term interest rates, we see that money expansion leads to decreases in the interest rate level. Moreover, similar to (Bjørnland and Leitemo 2009) the impact on industrial production is ambiguous with changing signs during the first five months after the shock. After that we see a standard result of money-neutrality in the long-run. As mentioned before, one can take the Mundell-Fleming model under a fixed exchange rate and with low capital mobility as point of comparison for China. The Mundell-Fleming model predicts for this case a decline in international reserves because the central bank uses reserves to support a specific exchange rate. This process is also called sterilization. In fact, this description is consistent with the empirical evidence from both models: international reserves fall and gradually rise over time again. An increase in domestic credit would - for an open economy with a floating exchange rate - tend to lead to a depreciation of the local currency and would rise the price of the foreign currency. If the central bank starts selling parts of their international reserves the foreign currency will relatively depreciate and in theory one can maintain the fixed exchange rate - at least as long as the central bank does not run out of international reserves. In contrast to model 1, the effect on international reserves is persistently negative for model 2.

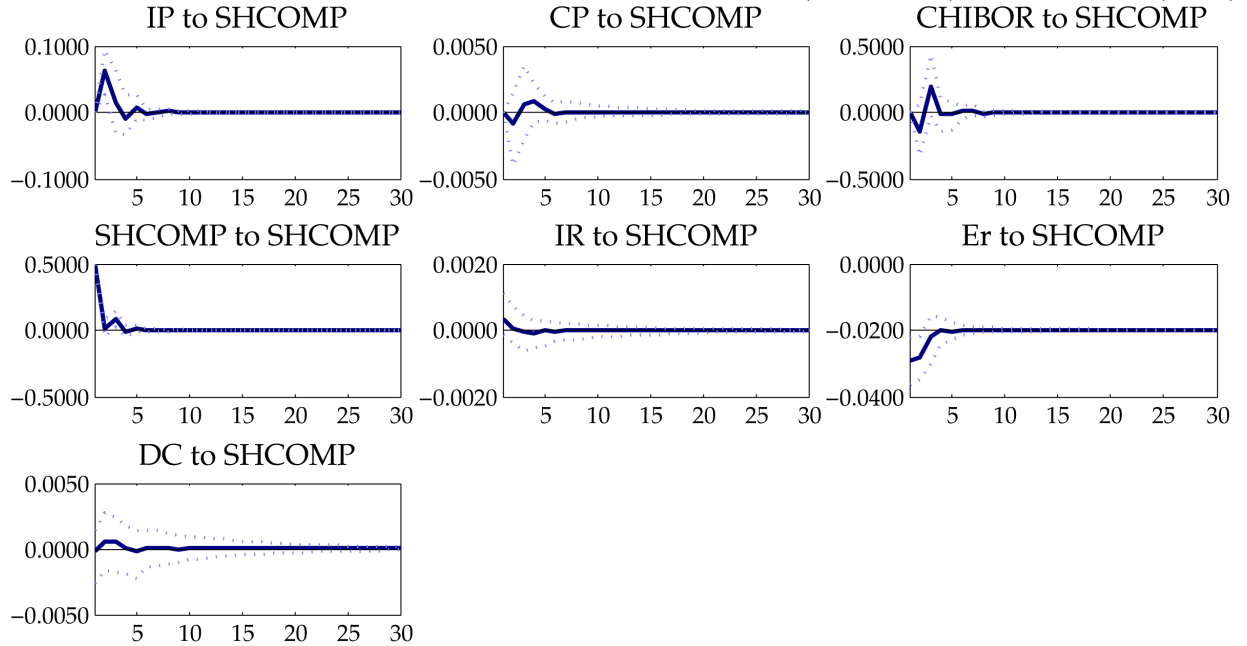
The effect on the effective real exchange rate shows some evidence of oscillations shortly after the shock occurs. This behavior distinguishes from the standard ‘overshooting’ in the Mundell-Fleming-Obstfeld model. Hereby, it is observable that the oscillations are stronger in

model 1. Moreover, increases in domestic credit lead to a positive short-run effect on the stock price index SHCOMP which is an intuitive result.

Theoretically, domestic credit growth has a liquidity effect - lowering real interest rates - and an expectations effect increasing interest rates. The latter typically implies that market participants expect higher inflation as a response to an increase in money supply and are therefore demanding higher interest rates to get compensated for the higher inflation. The empirical evidence for China suggests a persistent negative impact of shocks to domestic credit on the Interbank lending rates which is highest for the second model including stock prices. This result is remarkable as it suggests that the liquidity effect seems to outweigh the expectation effect even in the long-run for China. Typically, the liquidity effect only outweighs the expectation effect in the short-run as described for example by (Cochrane 1989).

## 6.4 Shocks to stock prices

Figure 10: Model 2 Impulse responses to unit stock price (SHCOMP) shock, 2 lags (in %)



In the short run a stock price shock increases Industrial Production Index growth. This can be theoretically explained by a wealth effect on consumption and a Tobin Q (Brainard and James 1968) effect on investment. However, the effect is not clear on consumer prices which is opposed to the clear positive effect on both variables in (Bjørnland and Leitemo 2009).

After about 2 months, there is a slight positive effect on the CHIBOR which in the context

of the positive impulse on Industrial Production makes intuitively sense. However, this effect is by far less severe than in (Bjørnland and Leitemo 2009) who found an interest rate increase of about 4 % as a response to a 1 % stock price shock. One might suppose that the reason for this lies in the difference between the federal funds rate considered in (Bjørnland and Leitemo 2009) and the Interbank lending rate (CHIBOR) considered here.

## 6.5 Robustness checks

As a standard practice in time series analysis several robustness checks are conducted. First of all, different levels of lags are considered. The Akaike information criterion suggested 2 lags and the impulse responses are calculated for variations of the number of lags. There are however no significant changes in the impulse responses, for example if the model is estimated with four lags.<sup>16</sup>

Replacing domestic credit by money supply provokes a diminished positive effect of a CHIBOR shock on money supply. However, as described before domestic credit seems to be a stronger proxy for monetary policy in this setting as it better captures open market operations conducted by the central bank. As an example the responses to CHIBOR shocks are included in figure 11 and in figure 12 which are relegated to the Appendix. In both models the impact of CHIBOR shocks on money supply is now negative, weakening the results from before.

The second specification option for these models based on the stationarity tests could have been a measurement of relative changes in the variables. The main impulse responses from domestic credit on CHIBOR shocks remain the same. As an example we include the impulse response to CHIBOR shocks in figure 13 and in figure 13 in the Appendix. Counterintuitive results, can be observed for model 1 when the CHIBOR rate increases as response to increases in domestic credit. The evidence on a wealth effect on consumption and a Tobin Q (Brainard and James 1968) effect on investment after a positive shock to stock prices is slightly stronger for this model specification.

However, taking model 2 and excluding CHIBOR also leads to counterintuitive results with stock prices declining as impulse response to positive industrial production shocks. The respective figure is also relegated to the Appendix in figure 15. In conclusion, empirical results

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<sup>16</sup>Details regarding the output of the robustness check are relegated to the Appendix.

support model 2 in comparison with both models that take into account variables in a different measurement form and in comparison with a model that excludes CHIBOR.

## 7 Conclusion

The empirical evidence suggests that Interbank lending rates interact with stock prices which have an impact on monetary policy decisions. This supports the results obtained by (Bjørnland and Leitemo 2009) and by (Bernanke and Kuttner 2005) for the US. Even for a country with low capital mobility stock markets seem to be essential to measure the interaction with monetary policy. Hence this gives reason for a model specification including stock prices such as demonstrated in model 2 rather than model 1 which only includes Interbank lending and omits stock prices.

For both model specification there is some empirical evidence that domestic credit as measure of monetary policy increases as response to rising CHIBOR rates. This meets the intuition that central banks might tend to inject money through open market operations into the market to push down Interbank lending rates. The positive impulse response of domestic credit to a CHIBOR shock is statistically significant for the first model. This substantiates the observation that as a response to increasing Interbank lending rates, central banks often follow an expansionary monetary policy.

On the other hand, the second model which includes stock prices shows a relatively better performance in accordance with conventional predictions by open market economy models. After including stock prices in the second model, the impulse response of domestic credit to a CHIBOR shock is still positive but not statistically significant at the 5 % level. Decreasing international reserves as response to rises in money supply are predicted by the Mundell-Fleming model for a country with fixed exchange rates.

All in all, one should prefer model 2 because stock prices play an important role in interaction with monetary policy. The results obtained for positive stock price shocks accord with (Bjørnland and Leitemo 2009) and reflect economic intuition. However, the impact of CHIBOR shocks on domestic credit is ambiguous and also not robust to other proxies of monetary policy. Nevertheless, model 2 seems favorable for example to a model excluding CHIBOR rate. In fact, leaving out CHIBOR leads to a negative impact of industrial production shocks on stock prices

which is counterintuitive. Therefore, it seems favorable to include Interbank lending rates as an additional variable.

Obviously, the empirical evidence is far from giving a clear response to the question if increases in CHIBOR rates in fact lead to increases in domestic credit. In the future, it might be beneficial to develop a theoretical model that better captures exactly this channel. Moreover, further research could rather focus on either an empirical investigation in other countries or be based on a model specification that takes into account the variables in a higher frequency. In particular, the latter could better capture the high dynamic of short-term rises in Interbank lending rates and measures taken by the central bank - namely expansion of the monetary base - as a response to these increases.

## 8 Appendix

### 8.1 Other model specifications

Figure 11: Model 1 Impulse responses to unit CHIBOR shock in model with money supply, 3 lags (in %)

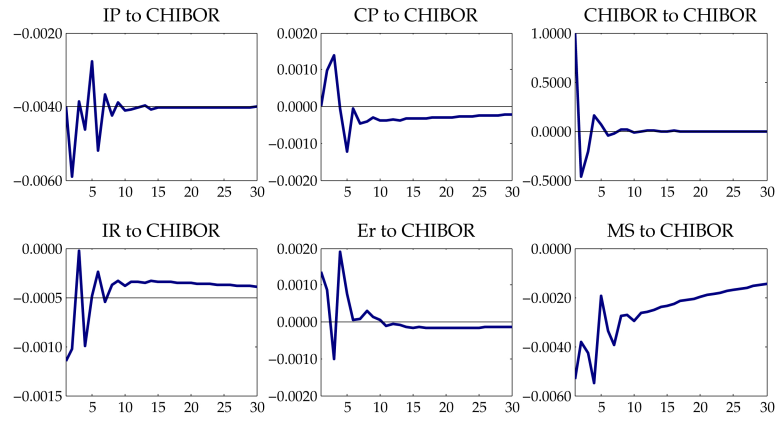


Figure 12: Model 2 Impulse responses to unit CHIBOR shock in model with money supply, 2 lags (in %)

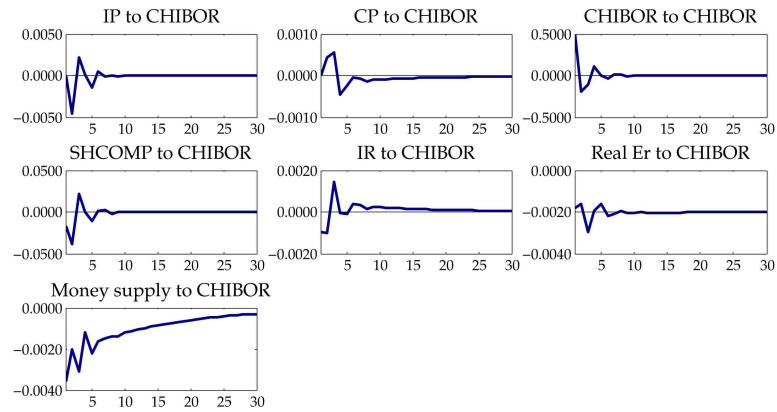


Figure 13: Model 1 Impulse responses to unit CHIBOR shock in relative model specification, 3 lags (in %)

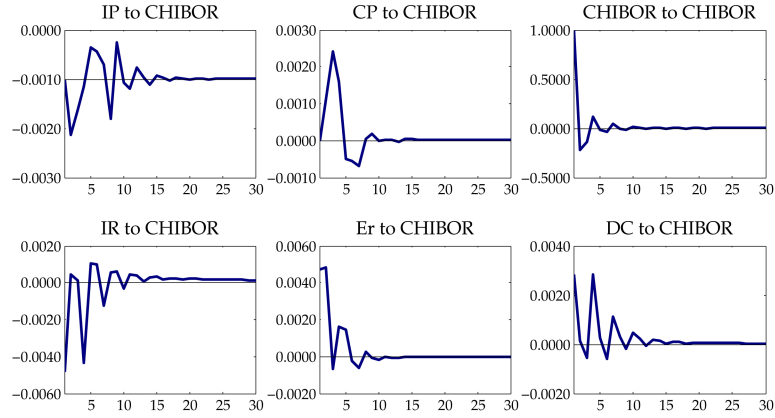


Figure 14: Model 2 Impulse responses to unit CHIBOR shock in relative model specification, 2 lags (in %)

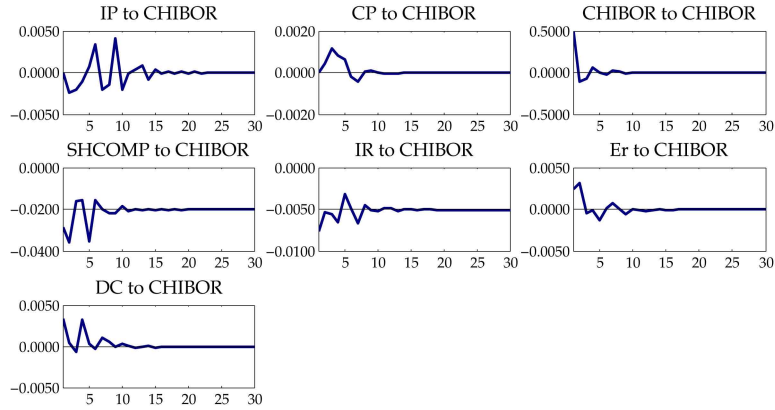


Figure 15: Model 2 excluding CHIBOR, impulse responses to unit Industrial Production shock, 2 lags (in %)

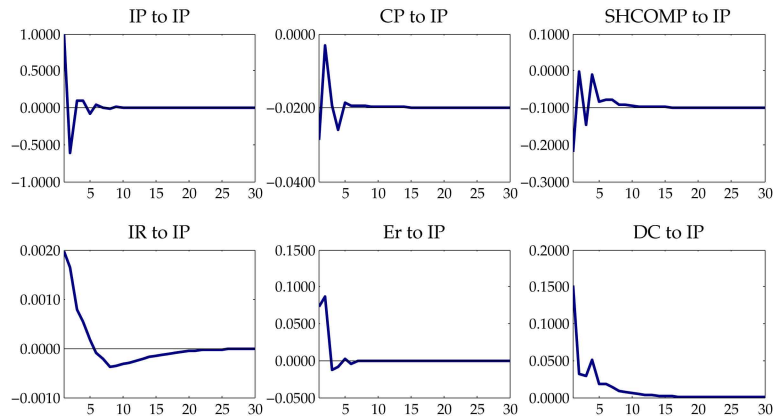
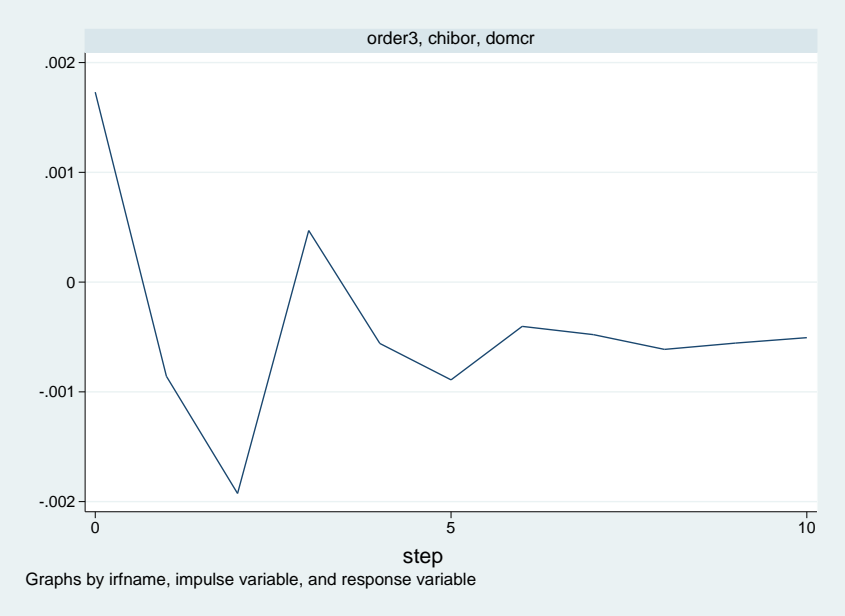


Figure 16: Model 2 as VECM, domestic credit impulse responses after CHIBOR shock, 2 lags (in %)



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