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Is Credit Easing Viable in Emerging and Developing Economies? An Empirical Approach

by Luis I. Jácome H., Tahsin Saadi Sedik, and Alexander Ziegenbein

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Abstract

During the global financial crisis, many central banks in advanced economies engaged in credit easing. These policies have been perceived as largely successful in reducing stress in financial markets, thus avoiding larger output losses. In this paper, we study empirically whether credit easing is also a viable policy tool to cope with banking crises in emerging and developing economies. We find that credit easing leads to a sharp increase in domestic currency depreciation, high inflation, and a substantial reduction in economic growth in a large panel of emerging and developing economies. For advanced economies, we find the effects to be benign. Our results suggest that emerging and developing economies should be cautious when using credit easing as it may fuel adverse macroeconomic repercussions.

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Keywords: Banking crises, credit easing, currency crises, financial stability, macroeconomic stability.

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Contents Page

Figures

Appendixes

I. INTRODUCTION

During the global financial crisis, many central banks in advanced economies engaged in credit easing.[2] Initially, these policies aimed to stabilize the financial system but were later also used to provide additional accommodation at the zero lower bound. Credit easing has been perceived as largely successful, especially at the time of greatest financial turmoil, when it averted the collapse of financial markets and thus prevented higher output losses. [3]

Credit easing has also been used in emerging and developing economies to mitigate financial stress in the banking system. Jácome (2008) documents the use of credit easing in Latin America during the 1990s and 2000s, whereas Ishi and others (2009), more broadly, illustrate how emerging market economies resorted to unconventional monetary policy, although in a limited amount during 2007–2009.

An important question is whether credit easing is also a suitable policy to deal with banking crises in emerging and developing economies. Specifically, can central banks in these countries draw on credit easing to cope with systemic banking crises without risking major macroeconomic repercussions? This question has received little attention in the literature. Theoretical models by Velasco (1987) and Chang and Velasco (1998) suggest that using large amounts of central bank money to ease conditions in the financial system may pave the way for a currency crisis. Finding an answer is crucial considering that emerging markets have a record of recurrent financial crises that took a large toll on economic growth and stability.

In this paper, we study empirically the effects of credit easing on key macroeconomic variables in a large panel of emerging and developing economies. We find that credit easing leads to a sharp currency depreciation, high inflation and a substantial reduction in economic growth. Our results suggest that emerging and developing economies should be cautious when using credit easing.

We use data on central bank credit to the financial system together with a recursive identification scheme to identify exogenous changes in credit easing. To estimate dynamic responses to the identified credit easing *shocks*, we use Jordà's (2005) local projection method. We first study the effects of credit easing in a linear a model and find that it leads to an increase in domestic currency depreciation, higher inflation and a reduction in growth. We then estimate a state-dependent local projection model allowing for the effects of credit easing to depend on whether the economy is experiencing a systemic banking crisis. Our results suggest that credit easing has large adverse side effects when implemented during an acute banking crisis—much larger than the linear model suggests. Our results are therefore consistent with the theoretical notions of Velasco (1987) and Chang and Velasco (1998). As a benchmark, we also study credit easing in a panel of advanced economies and find the

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[2] Bernanke (2009) distinguishes three broad categories of credit easing: (1) lending to financial institutions; (2) providing liquidity to key credit markets; and (3) purchasing longer-term securities.

[3] Joyce and others (2012) and Borio and Zabai (2016) summarize the existing literature on the effectiveness of credit easing and other unconventional monetary policies.

effects to be much more benign: credit easing has a mild positive effect on inflation and a positive but negligible effect on output growth.

Empirical evidence on the effects of credit easing in emerging and developing economies is scarce. Jácome and others (2012) find a significant relationship between central bank monetization and large currency depreciations in a sample of banking crises in Latin America. We expand on their work along several dimensions. First, we consider a much larger sample of emerging and developing countries, with broader geographical coverage. Second, we do not restrict ourselves to the impact of credit easing on exchange rates. Instead, we explore the dynamic effects of credit easing on other key macroeconomic variables, like inflation and output. Third, we differentiate between effects during rather calm times and systemic banking crises, making use of the flexibility of the local projection method. Finally, to establish a meaningful benchmark, we estimate the effects of credit easing in a panel of advanced economies. In addition, Kaminsky (1999) and Laeven and Valencia (2013) document that banking crises often precede currency crises. However, they do not identify central bank monetization as a mechanism connecting banking and currency crises as we do in this paper.

The rest of the paper is structured as follows. Section II describes the data and introduces the methodology. Section III discusses the empirical results. Section IV presents robustness checks. Section V explores extensions to the baseline specification. Section VI offers concluding remarks.

II. DATA AND METHODOLOGY

This section provides a detailed account of how we construct our panel data set and the econometric specification used in our analysis.

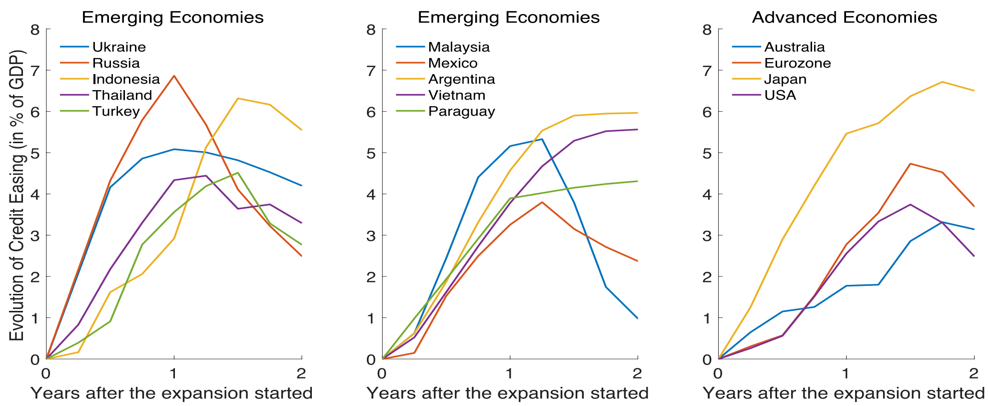
A. Data

We follow Jácome and others (2012) and Laeven and Valencia (2013), and use the central bank's claims on deposit money banks as well as claims on other financial institutions as a measure of credit easing. Our measure captures credit easing as defined by Bernanke (2009), as well as variations in central bank credit stemming from open market operations. It thus summarizes central bank's liquidity provision to the banking system.

In emerging and developing economies, the central bank extends credit to financial institutions for several purposes, including to help banks withstand deposit withdrawals and to facilitate bank resolution. This is similar to the Fed's Term Auction Facility introduced since August 2007, the European Central Bank's provision of credit in 2011 and 2012 to support banks in the peripheral countries facing deposit withdrawals, and the Bank of England's financial support to Northern Rock and, generally, its rescue plan to financial institutions in 2008.

To provide an idea of the size of credit easing in emerging markets, and compare it to the recent experience in advanced economies, Figure 1 shows the evolution of credit easing during major credit expansions for 10 emerging economies, three advanced economies and the Eurozone. To obtain a comparable metric across countries, we scale our variable by each country's nominal GDP. The start values are normalized to zero to facilitate such comparison.

Figure 1. Major Expansions of Credit Easing in Advanced and Emerging Economies



Sources: IFS, WEO, and authors' calculations.

Note: Credit easing is measured by central bank claims on banks and other financial institutions and scaled by nominal GDP.

To conduct the empirical analysis, we create a panel of emerging and developing economies, taking as a reference point the list of 145 countries that are classified by the IMF's *World Economic Outlook* (WEO) 2016 as emerging and developing economies. We then drop 49 countries that are listed as "least developed" by the United Nations for our econometric analysis due to concerns about data quality and availability, which leaves us with 96 countries. Since for 22 of these countries, we do not have sufficient data to perform our dynamic analysis, we end up with a panel of 74 emerging and developing economies. The list of the countries used in our analysis is presented in Appendix I.

To identify periods of banking crises, we use of the extensive database compiled by Laeven and Valencia (2013). We use quarterly data from 1995:Q1 to 2012:Q4. In total, the sample consists of 4,656 country-quarters. It includes 35 distinct systemic banking crisis episodes, and the total number of observations during banking crises is equal to 330 quarters. When quarterly data are not available, we convert annual frequency to quarterly frequency.[4] A detailed list of the variables used in the analysis with the corresponding sources and relevant explanations is presented in Appendix II.

As a benchmark, we compare the results for emerging and developing economies to those obtained for advanced economies. For this purpose, we build a separate panel consisting of

[4] To make sure that our results are not influenced by the interpolation method, we perform a number of robustness checks. First, we use three alternative methods for frequency conversion and run our analysis for each case. We use cubic splines, Litterman, or Denton frequency conversion. All results reported in this paper are robust to using any of these frequency conversions. In the figures shown in this paper, we use the Denton method. Second, we perform our analysis on the subset of countries for which we have sufficient quarterly data. Our findings remain unchanged.

13 advanced economies for the same period and use data from the same sources. We proceed in the same manner as with the emerging and developing economies sample. That is, we use the list of advanced economies/currency areas as classified in the WEO 2016, and use all economies for which we have sufficient data to perform our analysis. In our sample, we treat the Eurozone as one economy. We again use quarterly data from 1995–2012. In total, the sample consists of 858 country-quarters. It includes 10 distinct banking crisis episodes, and the total number of observations during banking crises is equal to 151 quarters. The list of advanced economies in the sample is also presented in Appendix I.

B. Methodology

We use local projections (LP; Jordà 2005) to estimate impulse responses to credit easing shocks.[5] We prefer LP to the more conventional vector autoregressions (VAR) for three reasons. First, the LP approach is more robust to misspecification and allows us to remain more agnostic about the underlying data generating process. Second, the LP approach is more flexible: it can easily be adapted to allow for state-dependent effects, while the computation of state-dependent impulse responses from a VAR is complex.[6] Third, when we use short run (or timing) restrictions, we can estimate structural impulse responses equation by equation using LP—one only needs to choose the appropriate set of controls.[7] With VARs, on the other hand, one needs to estimate the entire dynamic system. This feature makes the estimation of structural impulse responses more feasible in a large panel setting and greatly simplifies the construction of confidence bands.[8]

Our four key variables are real GDP growth (Y), inflation (π), credit easing (CE), and the nominal exchange rate (E). In our baseline specification, we add two control variables that we believe are important to pin down the causal effect of credit easing. First, we use the short-term interest rate[9] (r), since we are primarily interested in the effects of credit easing unrelated to open market operations. Second, we include an indicator variable (B) that takes

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[5] Recently an increasing number of studies have relied on local projections to estimate impulse response functions. Popular examples are Auerbach and Gorodnichenko (2011), Ramey and Zubairy (2014), and Romer and Romer (2015).

[6] In a regime-switching VAR, the researcher estimates a set of reduced form VAR parameters for each regime. The construction of impulse responses from these estimates requires making assumptions about how and when the parameters switch from one regime to the other. A common assumption is to impose that the shock of interest cannot alter the state within the impulse response horizon (see, for instance, Auerbach and Gorodnichenko 2012). This assumption, however, seems implausible in the context of credit easing shocks. We can sidestep this difficulty with the LP approach because the construction of impulse responses does not rely on the Wold decomposition theorem. With LP, we instead estimate the average effect of a shock that hits in a given state, without needing to specify further whether the shock causes the economy to transition from state to state.

[7] We provide formal proof of this in Appendix IV.

[8] With VARs, one has to rely on bootstrapping or the delta method. With LP, we estimate impulse responses equation by equation using ordinary least squares (OLS). Hence, we can use heteroskedasticity and autocorrelation (HAC) standard errors to construct confidence bands (Jordà 2005).

[9] We use data on the central bank policy rate where applicable. In economies that do not have an official policy rate, we use short-term deposit rates as a substitute. We explore alternative substitutes such as discount rates and money market rates and find that our results remain unchanged.

the value 1 if a banking crisis is ongoing, and 0 otherwise. As we have seen in the previous section, large changes in credit easing occur mostly during systemic banking crises. We add B to help isolate the effects of a banking crisis.

To identify structural impulse responses, we make an additional assumption. We adapt an assumption frequently used in the study of monetary policy in open economies. That is, we use the short-run restriction that macroeconomic variables react with a lag to monetary policy, while the nominal exchange rate reacts contemporaneously (see, for example, Cushman and Zha 1997, Kim and Roubini 2000, and Cologni and Manera 2008).

Therefore, our system of variables (X) is

$$X = [\, Y, \pi, \mathrm{B}, \mathrm{r}, \mathrm{CE}, \mathrm{E}]$$

C. Linear Model

To implement the timing restriction in LP, we have to include the contemporaneous values of variables that react with a lag to changes in credit easing as controls along with the lagged values of all variables. In short, we estimate a sequence of projections for each impulse response horizon h and each dependent variable of interest x:

$$(1)\ x_{i,t+h} = \sum_{j=0}^{L} \left(a_j^h Y_{i,t-j} + b_j^h \pi_{i,t-j} + c_j^h B_{i,t-j} + d_j^h CE_{i,t-j} + e_j^h r_{i,t-j} + f_j^h E_{i,t-j} \right) + \mu_i^h + \lambda_t^h + \epsilon_{i,t+h},$$

$$\text{for } h = 0, 1, 2 \ldots, H.$$

To implement our identifying assumption, we impose $f_0^h = 0$.[10] μ_i^h are country fixed effects, and λ_t^h are time fixed effects. The impulse response of variable x to a credit easing shock is then given by $IR(h) = d_0^h$. To study the dynamic effects of credit easing shocks, we consider a horizon of 10 quarters (H = 10), and we use two lags on all variables (L = 2).[11]

Our specification implies that a credit easing shock is orthogonal to contemporaneous values of output growth, inflation, the banking crisis indicator and the short-term interest rate as well as to lagged values of all variables. Thus, the credit easing shock is orthogonal to current economic conditions. Since the credit easing shock is orthogonal to the current short-term interest rate, it captures the effect of an increase in central bank credit that is not due to open market operations.

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[10] In terms of shock identification, this is equivalent to estimating a VAR for $X = [\, Y, \pi, \mathrm{B}, \mathrm{r}, \mathrm{CE}, \mathrm{E}]$ and using a Cholesky decomposition to pin down the structural shocks to credit easing (CE). We provide formal proof of this in Appendix III.

[11] The Bayesian information criterion (BIC) and Akaike information criterion (AIC) suggest two and four lags, respectively. We use the more parsimonious model here. However, the results using four lags instead are similar.

Serial correlation may be induced in regressions when the lead dependent variable is introduced in the regression (horizon h > 0). Following Jordà (2005), we account for serial correlation by reporting Newey-West (1987) corrected standard errors.

D. State-Dependent Model

When we estimate the linear model (1), we obtain estimates of the average effects of credit easing over all states. However, injecting central bank money into the banking system might have profoundly different effects depending on the level of stress in the banking system. Intuitively, during a smaller banking crisis, liquidity injections might be helpful as long as they do not interfere with the central bank's main policy objective of achieving price stability. During times of severe banking crises, however, economic agents might expect that the central bank will abandon its primary policy goal and inject large amounts of liquidity into the financial system. This, in turn, might trigger a substantial domestic currency depreciation and inflation, and, in some cases, pave the way for a currency crisis (Velasco 1987). Therefore, a linear model is likely to underestimate the effects (in absolute value) of credit easing during a systemic banking crisis. To address this issue, we estimate a state-dependent model:

$$(2)\ x_{i,t+h} = \sum_{j=0}^{L} B_{i,t-j}\left(a_j^{h,B} Y_{i,t-j} + b_j^{h,B}\pi_{i,t-j} + c_j^{h,B} B_{i,t-j} + d_j^{h,B} CE_{i,t-j} + e_j^{h,B} r_{i,t-j} + f_j^{h,B} E_{i,t-j}\right)$$
$$+\sum_{j=0}^{L}(1 - B_{i,t-j})\left(a_j^{h,N} Y_{i,t-j} + b_j^{h,N}\pi_{i,t-j} + c_j^{h,N} B_{i,t-j} + d_j^{h,N} CE_{i,t-j} + e_j^{h,N} r_{i,t-j} + f_j^{h,N} E_{i,t-j}\right)$$
$$+\mu_i^h + \lambda_t^h + \epsilon_{i,t+h}.$$

To implement our identifying assumption, we impose $f_0^{h,i} = 0$ for $i = \{B, N\}$. The impulse response of variable x to a credit easing shock during a systemic banking crisis is given by the coefficients $IR^B(h) = d_0^{h,B}$, and, during *normal* times, by $IR^N(h) = d_0^{h,N}$.

III. Results

We start by letting the data speak for themselves and thus plot our measure of credit easing together with three key variables (real GDP growth, inflation, and nominal exchange rate) for a large subset of countries. Four key observations emerge from these figures. First, large expansions of credit mostly occur during systemic banking crises. Second, credit easing is also used during non-systemic banking crises. Third, large expansions of credit easing do not occur during *all* banking crises. Fourth, there is substantial co-movement between credit easing and real GDP growth, inflation, and the nominal exchange rate (see Appendix III).

We next present the results of the linear and the state-dependent models to allow the effects of the credit easing to depend on the level of stress in the banking system. While the goal of the paper is to assess the dynamic effects of credit easing on key macroeconomic variables in emerging and developing economies, we also estimate the effects of credit easing in advanced economies as a benchmark.

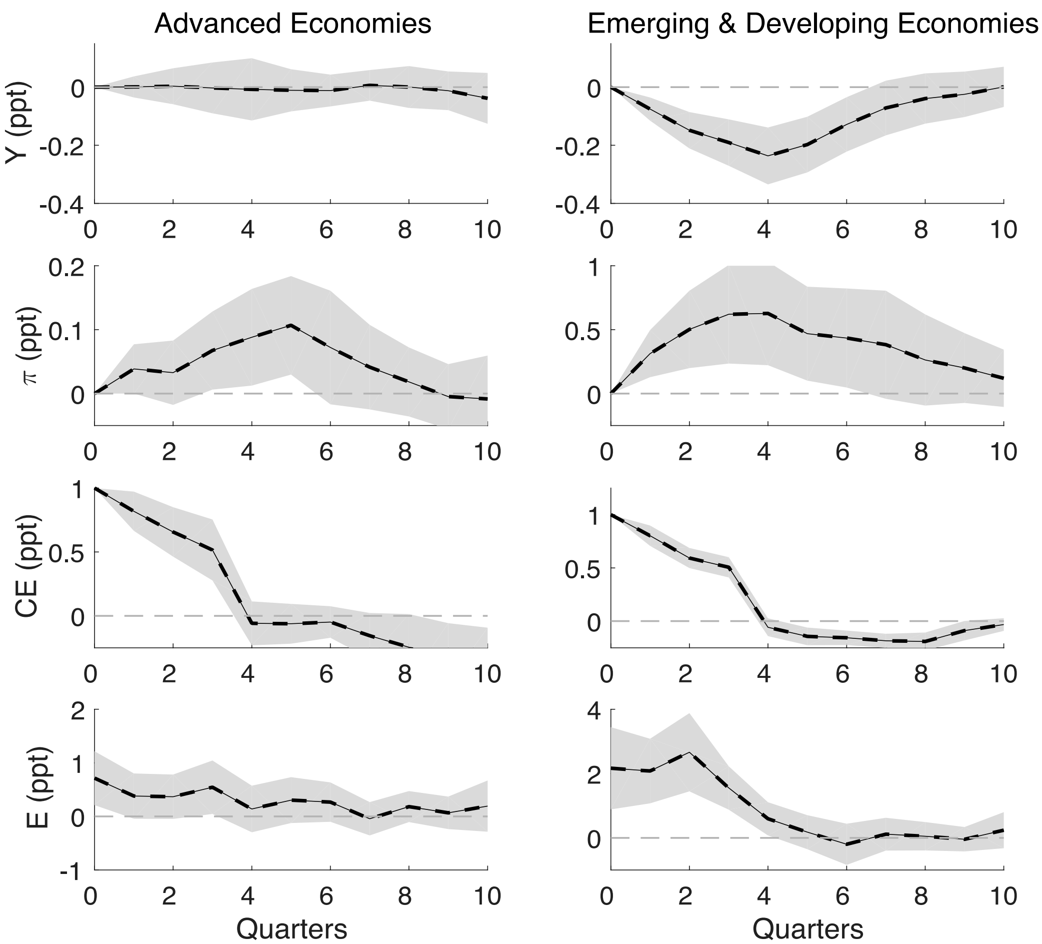
A. Linear Effects

We apply our econometric specification to the two panels separately and focus on the linear specification described in Equation (1). We estimate impulse responses to an increase in credit easing (equal to 1 percent of GDP) for real GDP growth, inflation, credit easing, and the nominal exchange rate. Figure 2 summarizes our results. The dashed black lines are point estimates for the impulse responses, and shaded areas are 90 percent confidence bands calculated using Newey-West standard errors.

In the left column, which shows the impulse responses for the sample of advanced economies, credit easing has the expected positive but moderate effect on inflation. Our estimates imply that the peak effect on inflation is roughly equal to 0.1 percentage points. Even if we consider an increase in credit easing, for instance of 5 percent of GDP, as in Figure 1, the peak effect on inflation would be equal to about 0.5 percentage points, the impact on output is negligible, and the exchange rate depreciates slightly.

The effects differ dramatically in the emerging and developing economies. An increase in credit easing equal to 1 percent of GDP leads to a peak effect on depreciation of about 2.7 percent and a peak effect on inflation of about 0.7 percent. The peak effect on inflation is seven times higher in emerging and developing economies than that in advanced economies. Moreover, we find that an increase in liquidity support hurts economic activity, reducing output growth by roughly 0.25 percent.

Figure 2. Impulse Responses to a Credit Easing Shock—Linear Model



Source: Authors' calculations.

Notes: Impulse responses of real GDP growth, inflation, credit easing, and nominal exchange rate depreciation to an increase in credit easing equal to 1 percent of GDP. Dashed lines are point estimates and shaded areas are 90 percent confidence bands calculated using Newey-West standard errors. Estimates from the linear local projection model using a panel of 13 advanced economies (left column) and a panel of 74 emerging and developing economies (right column) using quarterly data from 1995–2012.

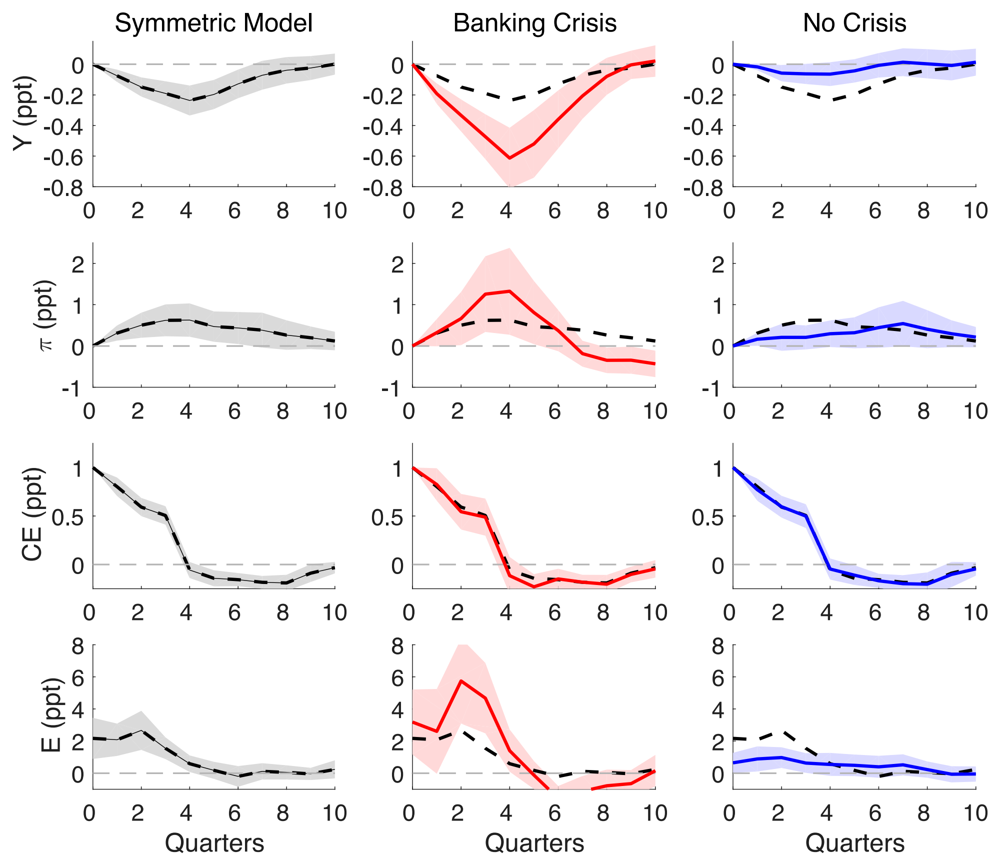
From a quantitative perspective, however, the results do not provide a clear answer on whether credit easing can be a link connecting banking and currency crises. The peak effects on growth, inflation, and depreciation might appear too small to assume that credit easing can trigger a currency crisis. However, we are here looking at the average effect of credit easing across all possible states of the economy. That is, our estimates provide an average of the effects of credit easing during times of minor or no disturbances in the banking system and the effects during systemic banking crises. We assess the macroeconomic effects of credit easing during systemic banking crises below.

B. State-Dependent Effects

To allow credit easing to have different effects depending on the level of stress in the banking system, we estimate the state-dependent model described in Equation (2) and, to identify systemic banking crises in emerging and developing economies, we use the Laeven Valencia's (2013) database.

Figure 3 summarizes our results. The left column shows the results of the linear model, which we considered in the previous section and serves as a benchmark. The dashed lines in all three panels are the impulse responses estimated from the linear model. Solid lines depict impulse responses estimated from the state-dependent model. The middle column shows the impulse responses to a credit easing shock equal to one percent of GDP during times of system banking crises. The right column shows the impulse responses to the same policy intervention during normal times. Again, shaded areas denote 90 percent confidence bands calculated using Newey-West standard errors.

Figure 3. Impulse Responses to a Credit Easing Shock—State-Dependent Model



Source: Authors' calculations.

Notes: Impulse responses of real GDP growth, inflation, credit easing, and nominal exchange rate depreciation to an increase in credit easing equal to 1 percent of GDP. Estimates from the linear model (dashed lines) and from the state-dependent model (solid lines) using a panel of 74 emerging and developing economies with quarterly data covering 1995–2012. The shaded areas are 90 percent confidence bands calculated using Newey-West standard errors.

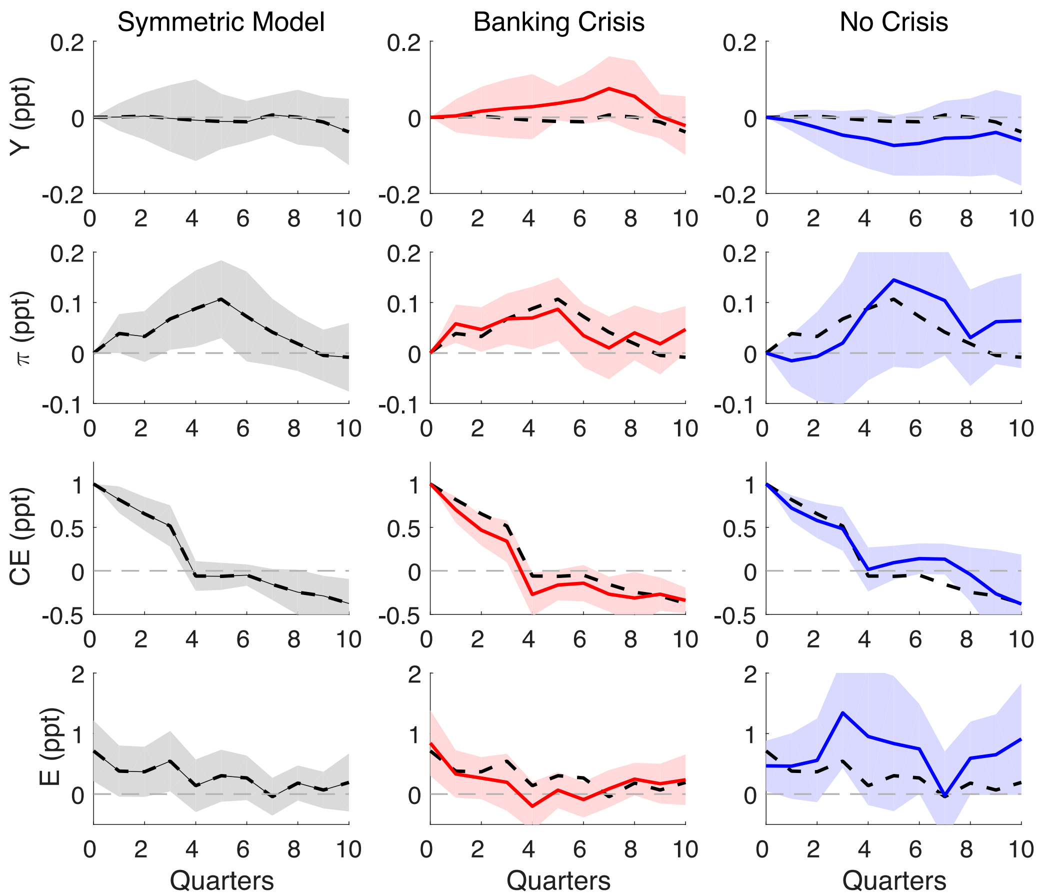
We find that the macroeconomic effects of credit easing differ dramatically across states. Credit easing during a systemic banking crisis (middle column) has much more extreme effects compared to the linear model (dashed lines). For a 1 percent of GDP credit easing shock, the peak effect on domestic currency depreciation is roughly 6 percent, on inflation is about 1.5 percent, and on output growth is -0.6 percent. In practice, expansionary credit policies have been considerably larger as illustrated in Figure 1 before. For instance, our results imply that an increase in credit easing equal to 5 percent of GDP leads to a currency depreciation of 30 percent, an increase in inflation of 7.5 percentage points, and a drop in economic growth of 3 percentage points. We also find that during normal times (right column) an expansion of central bank lending to financial institutions has a small or insignificant effect on the exchange rate, inflation, and output. Importantly, the effect on CE_t is roughly the same in both states, so we can rule out that our results are solely driven by more extreme or more persistent shocks during systemic banking crises. Our results indicate that accounting for state-dependent effects of credit easing is crucial to understanding the policy's dynamic effects on macroeconomic variables.

Our results for the state-dependent model are consistent with the notion that credit easing policies are a link connecting banking and currency crises. The effects of credit easing shock during a systemic banking crisis—a sharp depreciation of the domestic currency and a large and persistent reduction of output—are consistent with the existing evidence on the effects of currency crises (see, for instance, Cerra and Saxena 2008).

We also estimate the state-dependent model for the panel of advanced economies (Figure 4). We find that the effects of credit easing during a banking crisis are close to the ones estimated from the linear model. Credit easing has a small positive effect on currency depreciation and inflation and no significant effect on output growth during a banking crisis.

Comparing the middle columns of Figure 3 and Figure 4 reinforces our conclusion from the linear model: credit easing is a suitable policy tool for advanced economies aiming to boost inflation, while it bears considerable risks for emerging and developing economies. Our results suggest that emerging and developing economies should be cautious when using credit easing, as it is likely to fuel further macroeconomic unrest.

Figure 4. Impulse Responses to a Credit Easing Shock in Advanced Economies



Source: Authors' calculations.

Notes: Impulse responses of real GDP growth, inflation, credit easing, and nominal exchange rate depreciation to an increase in credit easing equal to 1 percent of GDP. Estimates from the linear model (dashed lines) and from the state-dependent model (solid lines) using a panel of 13 advanced economies with quarterly data covering 1995–2012. The shaded areas are 90 percent confidence bands calculated using Newey-West standard errors.

For advanced economies, our results are consistent with the large literature on the effects of credit easing and quantitative easing following the global financial crisis. The literature typically finds that these policies are effective in raising inflation, although the effect is short-lived.[12] There is no consensus about the effects on GDP. Our results are in line with previous studies that find small or insignificant effects on GDP (see, for instance, Schenkelberg and Watzka 2011; Chen, Curdia, and Ferrero 2012; and Pesaran and Smith 2014).

Why do the effects of credit easing in emerging and developing economies differ markedly from those in advanced economies? An important difference may be the trust in the domestic

[12] See Gagnon and others (2011), Krishnamurthy and Vissing-Jorgensen (2011), and Joyce and others (2012).

currency and trust in the central bank to maintain price stability. The main advanced economies, like the United States, the United Kingdom, Japan, and the Euro zone, issue reserve currencies, which are considered a “safe haven” in times of global financial turbulence. In these economies, extending loans to financial institutions—to avoid the collapse of financial markets—may not have a large impact on the exchange rate and inflation since central banks simply accommodate higher demand for liquidity. Moreover, in most advanced economies, central banks have a strong reputation and can credibly commit to maintaining price stability. This, in turn, keeps inflation expectations in check.

Emerging and developing economies not only do not issue reserve currencies, but many of them have a recent history of high inflation and currency crises that influences economic agents’ behavior. Thus, credit provision on a large scale may call into question the central bank’s role as a guarantor of price stability, giving way to an increase in expected inflation and currency depreciation. Moreover, many emerging markets and developing countries feature financial dollarization, which can make the financial system vulnerable to capital outflows and currency depreciation, as bank depositors withdraw their savings and protect them against inflation by shifting their assets to foreign currency. Under these conditions, if the effect on depreciation is large enough, this can lead to balance of payments problems and the large output losses associated with it.[13] Furthermore, for emerging and developing economies, even small currency depreciations are often contractionary events. If domestic firms and banks borrow in dollar, but equity is denominated in domestic currency, a depreciation, increases debt service payments and lowers the capacity to invest (see Serena and Sousa, 2017). Moreover, investment may decline further due to an *international borrowing constraint:* after a depreciation the collateral value of assets denominated in domestic currency falls and makes it more difficult to borrow from abroad (see Braggion, Christiano, and Roldos, 2009).

IV. Sensitivity Analysis

In this section, we study the robustness of our results. First, we repeat our analysis using an alternative crisis chronology. Second, we explore alternative identification assumptions. Third, we address worries over potential omitted variable bias by including additional control variables.

A. An Alternative Crisis Chronology

Until now, we have used the Laeven and Valencia (2013) database to identify systemic banking crises. However, there is more than one way of identifying banking crises and the chronology of crises may vary, as financial stress is worse in some periods than in others, making it often hard to draw the line between “crisis” and “no crisis.” Therefore, alternative crisis chronologies deviate from the Laeven and Valencia measure at times.[14] The most

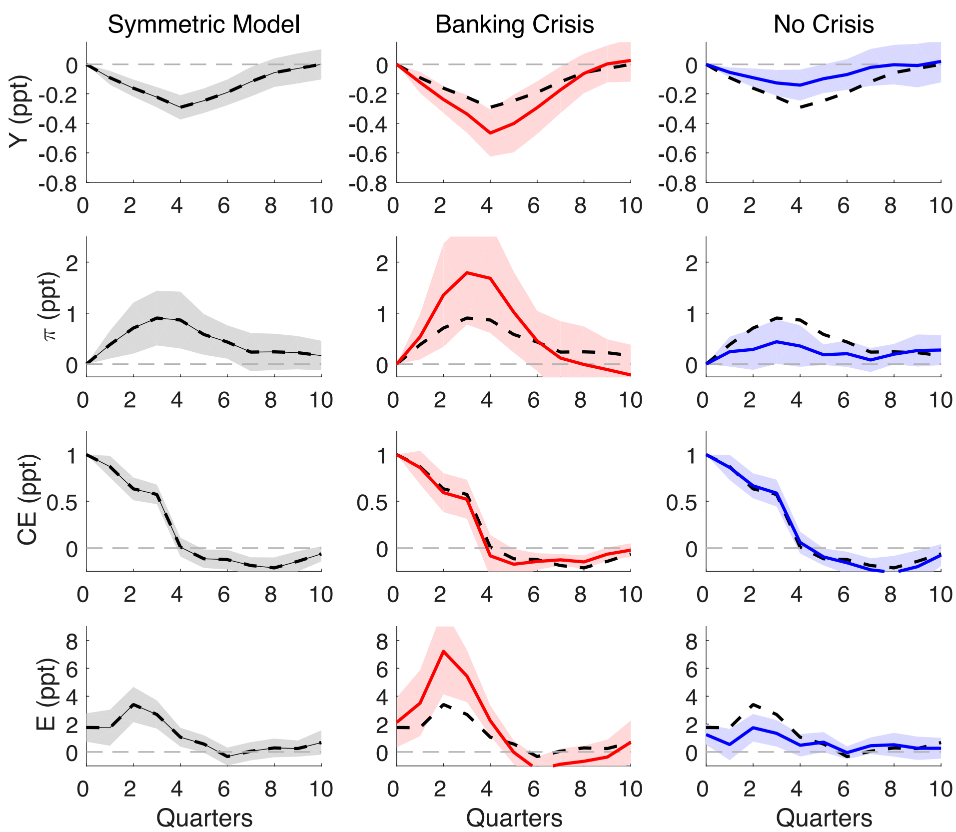
[13] Cerra and Saxena (2008) find that a balance-of-payment crisis lowers output by approximately 4 percent, on average.

[14] Romer and Romer (2015) gives a good description of deviations across crisis chronologies.

widely used banking crisis database aside from Laeven and Valencia is the one compiled by Reinhart and Rogoff (2009). In this section, we re-estimate Equation (1) and Equation (2) using the crisis chronology by Reinhart and Rogoff for our banking crisis indicator variable B.

Figure 5 summarizes our results. Comparing Figure 3 and Figure 4, we find that the results are broadly similar for the two alternative crisis chronologies. For the specification using the Reinhart and Rogoff chronology, we find somewhat larger peak effects of credit easing shocks on exchange rate depreciation and inflation, and a smaller effect on GDP growth during a systemic banking crisis.

Figure 5. Impulse Responses to a Credit Easing Shock—Alternative Crisis Chronology



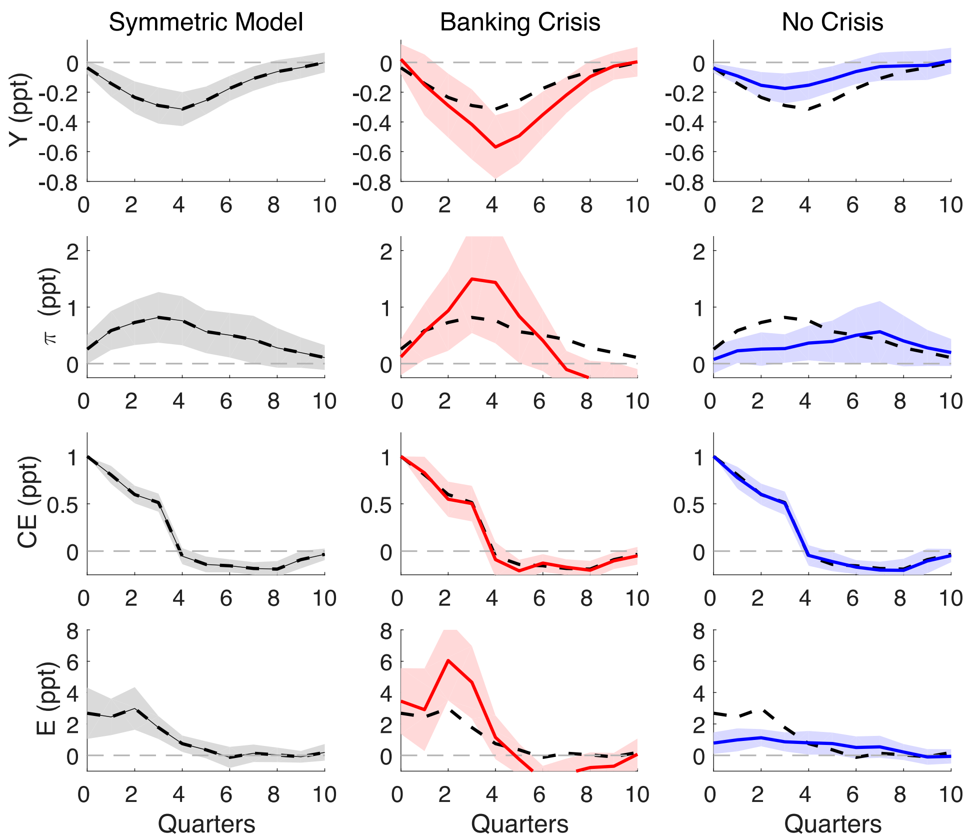
Source: Authors' calculations.

Notes: Impulse responses of real GDP growth, inflation, credit easing, and nominal exchange rate depreciation to an increase in credit easing equal to 1 percent of GDP. Estimates from the linear model (dashed lines) and from the sign-dependent model (solid lines) using a panel of 74 emerging and developing economies with quarterly data covering 1995–2008. The shaded areas are 90 percent confidence bands calculated using Newey-West standard errors.

B. Identification Assumption

We identify the effects of credit easing shocks by assuming that macroeconomic variables react with a lag to monetary policy while the nominal exchange rate reacts contemporaneously. While we find this assumption the most plausible and consistent with the literature, we explore other alternatives and apply them to our sample of emerging and developing countries.

Figure 6. Impulse Responses to a Credit Easing Shock—Alternative Identification Assumption I



Source: Authors' calculations.

Notes: Impulse responses of real GDP growth, inflation, credit easing, and nominal exchange rate depreciation to an increase credit easing equal to 1 percent of GDP. Credit easing is ordered first. Estimates from the linear model (dashed lines) and from the state-dependent model (solid lines) using a panel of 74 emerging and developing economies with quarterly data covering 1995–2012. The shaded areas are 90 percent confidence bands calculated using Newey-West standard errors.

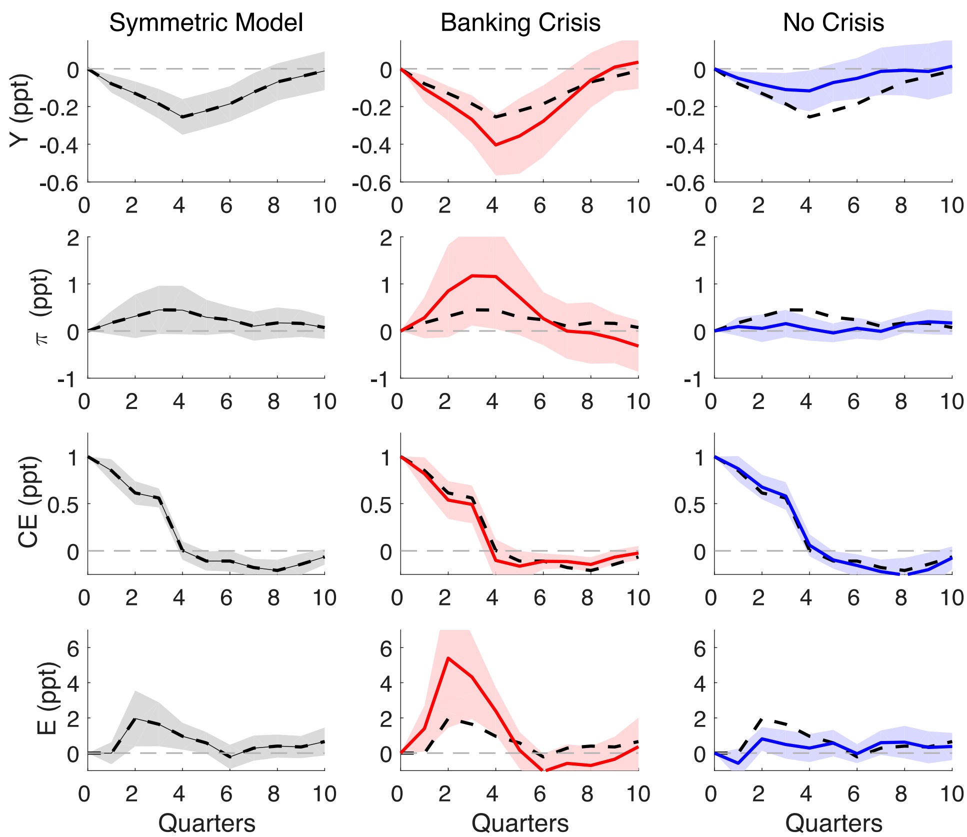
For illustration, we consider the two most extreme cases. First, we assume that *all* variables can react contemporaneously to changes in credit easing. To implement this assumption, we estimate Equation (2) and impose $a_0^{h,i} = b_0^{h,i} = c_0^{h,i} = d_0^{h,i}, f_0^{h,i} = 0$.[15] Figure 6 summarizes the results. Second, we assume the other extreme alternative, that is, none of the variables reacts contemporaneously to changes in credit easing. To implement this assumption, we

[15] In terms of shock identification, this is equivalent to estimating a VAR for $X = [CE, Y, \pi, B, r, E]$ and using a Cholesky decomposition to pin down the structural shocks to credit easing (CE).

(continued…)

estimate Equation (2) without additional restrictions on the parameters.[16] The results of this exercise are summarized in Figure 7. Comparing the results of our baseline specification in Figure 3 with Figure 6 and Figure 7, respectively, we see that the results are remarkably similar across all three specifications, which allows us to conclude that our results are robust to assuming alternative causal orderings of the variables.

Figure 7. Impulse Responses to a Credit Easing Shock—Alternative Identification Assumption II



Source: Authors' calculations.

Notes: Impulse responses of real GDP growth, inflation, credit easing, and nominal exchange rate depreciation to an increase credit easing equal to 1 percent of GDP. Credit easing is ordered last. Estimates from the linear model (dashed lines) and from the state-dependent model (solid lines) using a panel of 74 emerging and developing economies with quarterly data covering 1995–2012. The shaded areas are 90 percent confidence bands calculated using Newey-West standard errors.

C. Omitted Variables

To address potential concerns about an omitted variable bias, we explore whether our results are robust to the inclusion of additional control variables. First, we include growth in credit to the private sector. In our baseline specification, we argue that contemporaneous values of

[16] In terms of shock identification, this is equivalent to estimating a VAR for $X = [Y, \pi, \mathrm{B}, \mathrm{r}, \mathrm{E}, \mathrm{CE}]$ and using a Cholesky decomposition to pin down the structural shocks to credit easing (CE).

output growth, inflation, the crisis indicator the short-term interest rates as well as lagged values of all variables in the system describe well the dynamics of the economy. In particular, we argue that these variables can capture the dynamics in a crisis, thus assuring that the credit easing shock is orthogonal to current economic conditions. We now include credit growth to more precisely isolate the effects of stress in the banking system. Second, we control for the stance of fiscal policy and include primary deficits scaled by nominal GDP. Third, we control for countries' external financial strength by including international reserves scaled by nominal GDP.[17] Low international reserves may shake depositors confidence, in particular in countries that have a peg and/or are financially dollarized, which could lead to further exchange rate pressures. Fourth, we control for changes in terms of trade to take into account exchange rate movements due to changes in commodity prices (see Cashin and others 2002, and Chen and Rogoff 2002). Fifth, we control for the degree of legal central bank independence, as a measure of central bank credibility, using the index developed by Dincer and Eichengreen (2014). We seek to test to what extent credible central banks can better mitigate the uncertainty typically associated with events of financial stress and provide financial assistance to impaired banks without further damage in terms of output, exchange rate depreciation, and inflation.[18]

For simplicity, we augment Equation (1) and Equation (2) with both lagged and contemporaneous values of all the additional control variables in our regression.[19] Our estimates are summarized in Figure 8.[20] Comparing the results from our baseline specification in Figure 3 to Figure 8, we find that estimates are similar and conclude that our results are robust to the inclusion of additional control variables. The main difference is that the additional controls "mop up" the residual variance of the projections and thereby improve the precision of parameter estimates. Hence, confidence bands in Figure 8 are somewhat thinner.

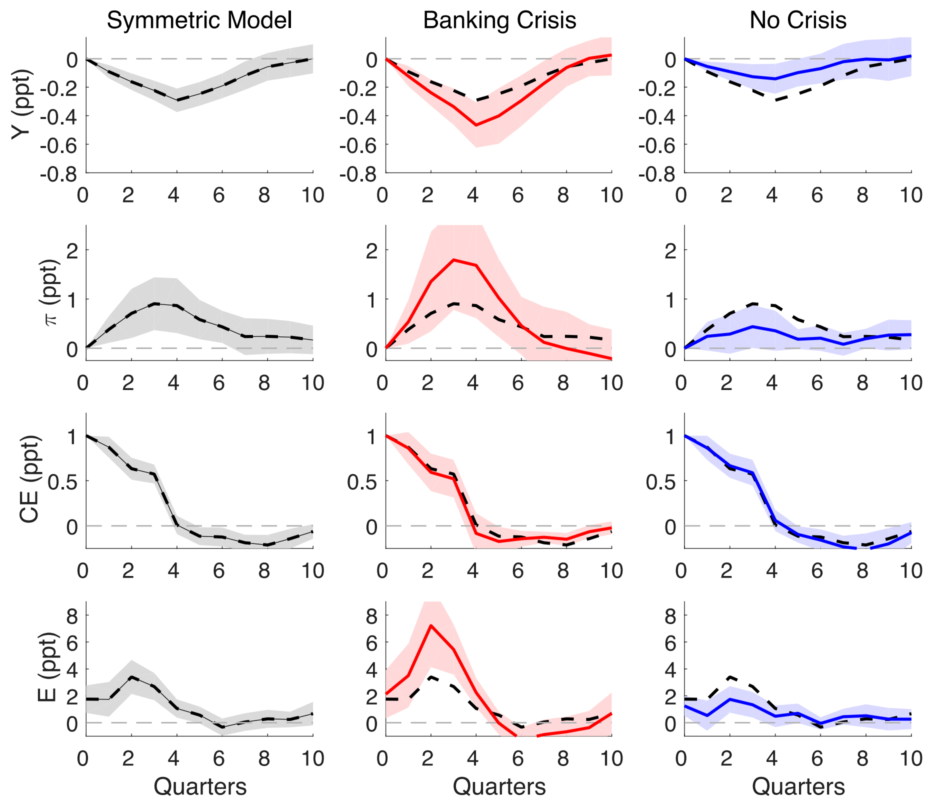
[17] We also scaled international reserves by broad money (M2). The results remain unchanged.

[18] We also ran the specifications using central bank independence or public deficits (or debt levels) as state-indicators. We find no significant asymmetries.

[19] Our sample is somewhat smaller, as data on central bank independence only cover 1998–2010.

[20] We repeat our previous exercise and check that the results are not sensitive to assuming that some control variables react contemporaneously to changes in credit easing. Our key findings remain unchanged.

Figure 8. Impulse Responses to a Credit Easing Shock—Additional Control Variables



Source: Authors' calculations.

Notes: Impulse responses of real GDP growth, inflation, credit easing, and nominal exchange rate depreciation to an increase credit easing equal to 1 percent of GDP. Estimates from the linear model (dashed lines) and from the state-dependent model (solid lines) using a panel of 74 emerging and developing economies with quarterly data covering 1998–2010. The shaded areas are 90 percent confidence bands calculated using Newey-West standard errors.

V. Extensions

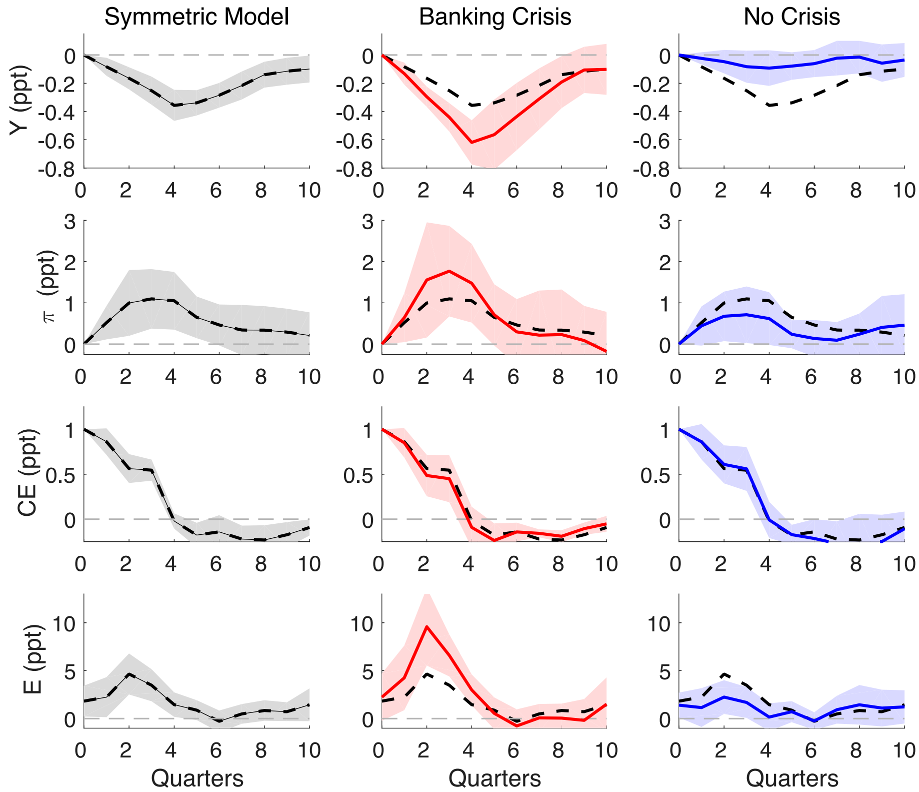
In this section, we consider two extensions to our baseline specification. We first split our sample of countries between emerging market economies and developing economies (excluding low-income countries), and study whether the effects of credit easing shocks differ across the two groups. Second, we study whether credit easing has sign-dependent effects, that is, whether expansions and contractions in credit policy have asymmetric effects.

A. Emerging Markets versus Developing Economies

Emerging market economies are more closely integrated into the global financial markets than developing economies. As a result, they are more prone to experiencing large capital outflows and currency depreciation than developing economies. Consequently, it is likely that the effects of credit easing are more pronounced in emerging market economies. To test this hypothesis, we split our sample into two groups—emerging market economies and developing economies—and study whether the dynamics unfolded by a credit easing shock are similar in those two groups.

As before, we follow the IMF's WEO 2016 to identify to which group a country belongs. Then, we estimate Equation (1) and Equation (2) for each group separately. The impulse responses for the sample emerging market economies are summarized in Figure 9. Figure 10 summarizes our results for the sample of developing economies.

Figure 9. Impulse Responses to a Credit Easing Shock—Emerging Market Economies

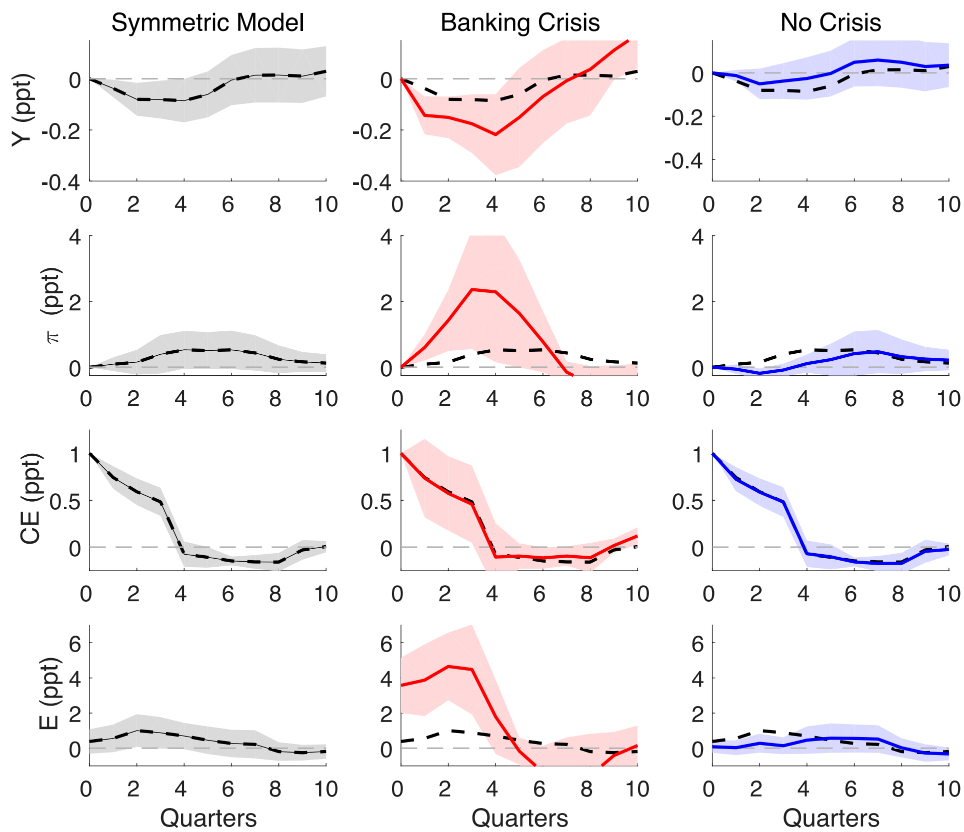


Source: Authors' calculations.

Notes: Impulse responses of real GDP growth, inflation, credit easing, and nominal exchange rate depreciation to an increase in credit easing equal to one percent of GDP. Estimates from the linear model (dashed lines) and from the state-dependent model (solid lines) using a panel of 20 emerging market economies with quarterly data covering 1995–2012. The shaded areas are 90 percent confidence bands calculated using Newey-West standard errors.

We find that the effects of credit easing are quantitatively larger in the panel of emerging market economies. However, the effects are qualitatively the same in both groups of countries. Therefore, our conclusions drawn in previous sections are valid for both groups of countries.

Figure 10. Impulse Responses to a Credit Easing Shock—Developing Economies



Source: Authors' calculations.

Notes: Impulse responses of real GDP growth, inflation, credit easing, and nominal exchange rate depreciation to an increase in credit easing equal to 1 percent of GDP. Estimates from the linear model (dashed lines) and from the state-dependent model (solid lines) using a panel of 54 developing economies with quarterly data covering 1995–2012. The shaded areas are 90 percent confidence bands calculated using Newey-West standard errors.

B. Sign-Dependent Effects

So far we have assumed that expansionary and contractionary credit policies have linear effects. However, there is no clear reason this should be the case. In fact, recent studies find significant sign-dependence in the effects of macroeconomic shocks.[21] The following scenario illustrates the importance of sign-dependent effects. Imagine a central bank expands credit to improve conditions in financial markets. As our previous results suggest, this might lead to a depreciation of the domestic currency, higher inflation, and a reduction in economic growth. As the central bank becomes aware of these adverse developments, it might want to revert its policy and engage in a contractionary credit policy. This, however, will only work if a contractionary credit policy has the mirror effect of an expansionary policy. Going back

[21] Barnichon and Matthes (2015a, 2015b) show that expansionary monetary and fiscal policy have a much smaller effect on output than contractionary policies. Barnichon, Matthes, and Ziegenbein (2016) find that sudden credit supply contractions have large effects on economic activity while credit supply expansions have little effect.

to the figures in Appendix III, we see that indeed such policy reversal often takes place while the systemic banking crisis is still ongoing. We can adapt our specification to formally test whether expansionary and contractionary credit policies have asymmetric effects. We now estimate:

$$(3)\ x_{i,t+h} = \sum_{j=0}^{L} I_{i,t-j}\left(a_j^{h,+}Y_{i,t-j} + b_j^{h,+}\pi_{i,t-j} + c_j^{h,+}B_{i,t-j} + d_j^{h,+}CE_{i,t-j} + e_j^{h,+}r_{i,t-j} + f_j^{h,+}E_{i,t-j}\right)$$
$$+\sum_{j=0}^{L}(1 - I_{i,t-j})\left(a_j^{h,-}Y_{i,t-j} + b_j^{h,-}\pi_{i,t-j} + c_j^{h,-}B_{i,t-j} + d_j^{h,-}CE_{i,t-j} + e_j^{h,-}r_{i,t-j} + f_j^{h,-}E_{i,t-j}\right)$$
$$+\mu_i^h + \lambda_t^h + \epsilon_{i,t+h}.$$

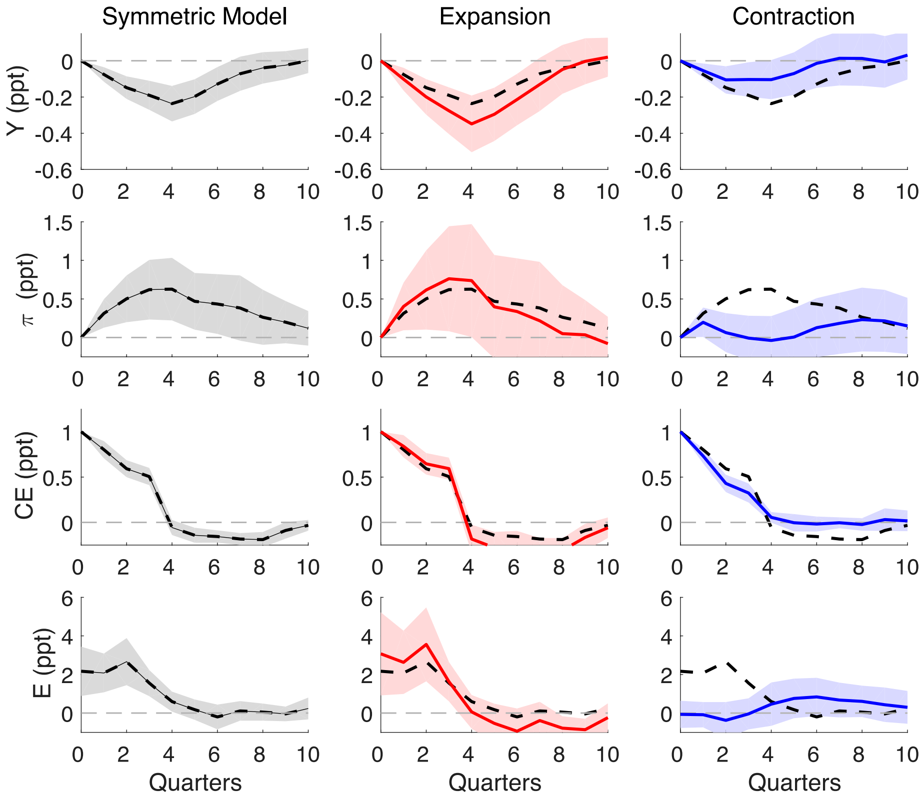
$I_{i,t}$ is an indicator variable that takes value 1 if $CE_{i,t} > 0$ and zero otherwise. To implement our identifying assumption, we impose $f_0^{h,i} = 0$ for $i = \{+,-\}$. The impulse response of variable x to an expansionary credit policy shock is given by the coefficients $IR^{+}(h) = d_0^{h,+}$, and to a contractionary credit policy shock by $IR^{-}(h) = d_0^{h,-}$.

Figure 11 summarizes our results. In the left column, we again see our results from the linear model as a benchmark. The dashed black lines are the point estimates of the linear model in all three panels. The solid lines are point estimates from the sign-dependent model, and shaded areas are 90 percent confidence bands calculated using Newey-West standard errors. The middle column shows impulse responses to an expansionary credit policy, that is, an increase in lending to financial institutions equal to 1 percent of GDP. The right column shows impulse responses to a contractionary credit policy. Impulse responses in the right column are multiplied by minus one for ease of comparison.

The results for sign-dependence are striking. We find that an expansionary credit policy has stronger effects than the linear model implies. Looking at the right column, we find that a contractionary credit policy has no significant effect on inflation or the nominal exchange rates and a small and short-lived effect on economic growth. Importantly, the effect on CE is roughly the same for expansionary and contractionary credit policies, so we can rule out that our results are driven by more extreme or more persistent expansionary shocks.

Our findings imply that while expansionary credit policy fuels further economic turmoil, contractionary credit policy has little effects. As such, our results suggest that once set into motion, the macroeconomic instability created by a credit expansion is not contained by reversing the policy. This insight should make policymakers in emerging and developing economies more cautious to engage in large credit expansions ex-ante.

Figure 11. Impulse Responses to a Credit Easing Shock—Sign-Dependent Model



Source: Authors' calculations.

Notes: Impulse responses of real GDP growth, inflation, credit easing, and nominal exchange rate depreciation to an increase in credit easing equal to 1 percent of GDP. Estimates from the linear model (dashed lines) and from the sign-dependent model (solid lines) using a panel of 74 emerging and developing economies with quarterly data covering 1995–2012. The shaded areas are 90 percent confidence bands calculated using Newey-West standard errors. Impulse responses to a contractionary shock (right column) are multiplied by minus one for ease of comparison.

VI. Conclusion

In this paper, we have studied the effects of credit easing on key macroeconomic variables in a large panel of emerging and developing economies. We found that credit easing leads to a depreciation of the domestic currency, inflation, and a reduction in economic growth. We also used a state-dependent model to allow the effects of credit easing to depend on the level of stress in the banking system. We found that the adverse side effects are dramatically higher during a systemic banking crisis.

We conclude from the empirical analysis that central banks in emerging and developing economies should be cautious when using credit easing to cope with financial crises. Our result also supports the idea that credit easing is an important link connecting banking and currency crises.

In addition, we have studied the existence of asymmetric (or sign-dependent) effects. We found that an expansionary credit policy has large effects on the exchange rate, inflation and output growth, while a contractionary policy has no significant effects. Our findings, therefore, suggest that the macroeconomic instability created through the balance sheet expansion cannot be contained by reversing the policy.

Our findings have important policy implications for emerging and developing economies. Given that credit easing can have large adverse side effects in the context of banking crises, countries should focus on reducing financial market vulnerabilities ex-ante by strengthening financial supervision and regulation and by establishing a strong macroprudential policy function. Stronger supervision and regulation will allow financial authorities to impose corrective actions before liquidity and capital shortages become severe. Macroprudential policy, in turn, should monitor and prevent the buildup of systemic vulnerabilities, which otherwise could end up in banking and currency crises; macroprudential policy can also increase resilience to shocks, including shocks associated with the reversal of capital flows. Yet, since crises are likely to occur inevitably, countries should put in place effective institutional underpinnings for bank restructuring and resolution. Otherwise, countries would inexorably resort to central bank money on a large scale, with the negative consequences discussed in this paper. Moreover, reducing currency mismatches beyond the banking system as well as financial dollarization can help to break the vicious cycle that connects banking and currency crises.

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Appendix I. Countries in the Sample

Emerging and Developing Economies	
Albania	Kuwait
Algeria	Kyrgyz Republic
Argentina*	Latvia
Armenia	Libya
Azerbaijan, Rep.	Lithuania
Barbados	Macedonia, FYR
Belize	Malaysia*
Bolivia	Mauritius
Botswana	Mexico*
Brazil*	Moldova
Brunei Darussalam	Mongolia
Cabo Verde	Morocco
Cameroon	Namibia
Chile*	Nicaragua
China, P.R.*	Nigeria
Colombia*	Pakistan*
Comoros	Papua New Guinea
Congo, Republic	Paraguay
Costa Rica	Peru*
Côte d'Ivoire	Philippines*
Dominica	Poland*
Dominican Republic	Romania*
Egypt	Russian Federation*
Fiji	Serbia, Republic
Gabon	Seychelles
Georgia	South Africa
Ghana	Sri Lanka
Grenada	Swaziland
Guatemala	Syrian Arab Republic
Guyana	Tajikistan
Honduras	Thailand*
Hungary*	Trinidad and Tobago
Indonesia*	Turkey*
Iran, I.R. of	Ukraine*
Jamaica	Uruguay
Jordan	Venezuela, Rep.*
Kenya	Vietnam

* Defined as EMEs in WEO 2016

Advanced Economies

Australia	Korea
Canada	New Zealand
Denmark	Norway
Euro Area	Sweden
Iceland	Switzerland
Israel	United States
Japan	

Appendix II. Description and Source of Variables

Variable	Description	Source
Nominal Exchange Rate	Value of the national currency per US dollar	IFS, line rf
Credit Easing	Central bank claims on deposit money banks + claims on other financial institutions divided by nominal GDP	IFS, line 12e + 12f
Real GDP	Gross domestic product at constant prices	WDI, WEO, IFS
Nominal GDP	Gross domestic product at current prices	WDI, WEO, IFS
Inflation	Annual rate of change in consumer prices	IFS, line 64
Short-term interest rate	Short-term deposit rate	IFS, line 60l
Policy rate	Central bank policy rate	IFS, line 60l
Private credit	Bank claims on private sector divided by nominal GDP	IFS, line 22d
International reserves	Total reserves minus gold divided by nominal GDP	IFS, line 1l.d
Public debt over GDP	General government gross debt divided by nominal GDP	WEO, series GGXWDGCD
Government Deficit	Central government net lending	WEO, series GGXCNL
Terms of Trade	Terms of trade in goods and services	WEO, series TT
Banking Crises I	Indicator variable that take value 1 during banking crises and 0 otherwise	Laeven and Valencia, 2013
Banking Crisis II	Indicator variable that take value 1 during banking crises and 0 otherwise	Reinhart and Rogoff, 2009
CBI	Central bank independence index	Dincer and Eichengreen, 2014

IFS: IMF International Financial Statistics
WEO: IMF *World Economic Outlook*
WDI: World Bank World Development Indicators
AREAER: IMF Annual Report on Exchange Arrangements and Exchange Restrictions

Data Adjustments:

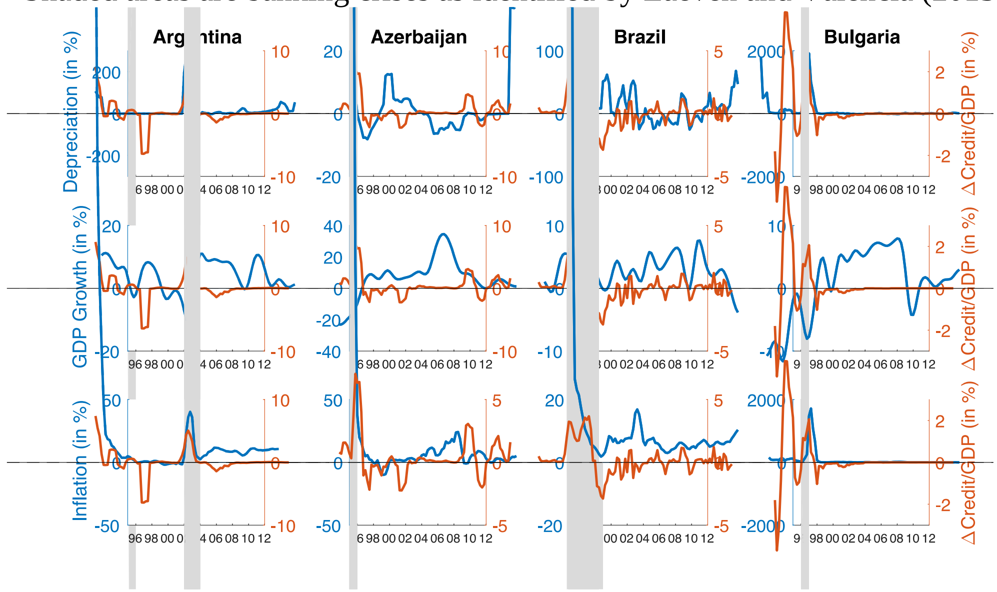
GDP data: In some instances, data on GDP (real and nominal) are not available at quarterly frequency. In those cases, we use annual data and use the proportional Denton method in Eviews12 to interpolate annual into quarterly frequency. To make sure our results are unaffected by the choice of the conversion method, we check the robustness of our results in two ways. First, we repeat our analysis, leaving out all countries that do not report quarterly GDP data. Second, we repeat our analysis using two alternative conversion methods: Litterman frequency conversion and cubic splines. Our results remain unchanged. Leaving out countries that do not report quarterly GDP data leads to slightly larger confidence bands due to the smaller amount of observations.

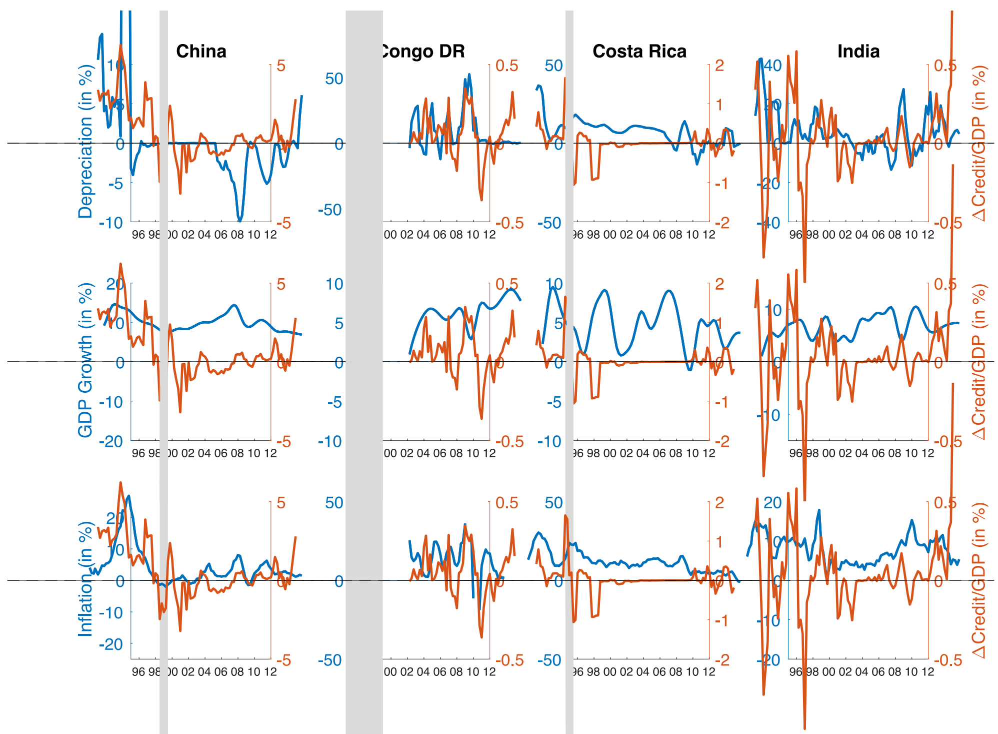
Terms of trade and public debt data: In most cases, data on terms of trade and general government debt are only available at annual frequency. We use the proportional Denton method in Eviews12 to interpolate annual into quarterly frequency. While we use both series only for one of our robustness checks and both can be considered slow moving, we again make sure that our choice of frequency conversion does not affect our results in a significant way. We perform the same steps as for the GDP data and find that our results remain unchanged.

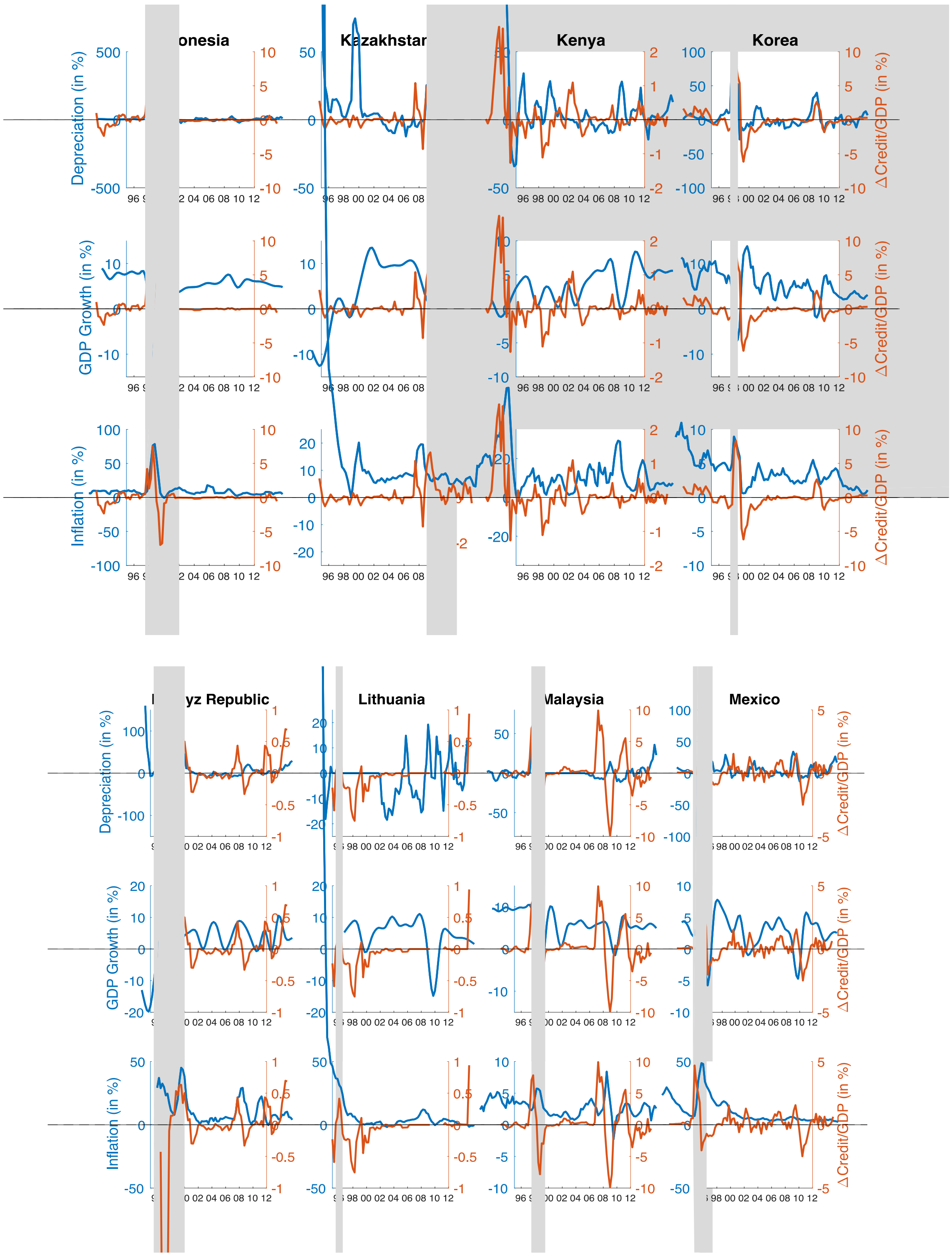
Banking crises: Laeven and Valencia (2013) report start and end dates of banking crises. We code our indicator variable such that it takes value 1 from start date through end date and 0 otherwise.

Appendix III. Central Bank Liquidity Support and Key Macroeconomic Variables

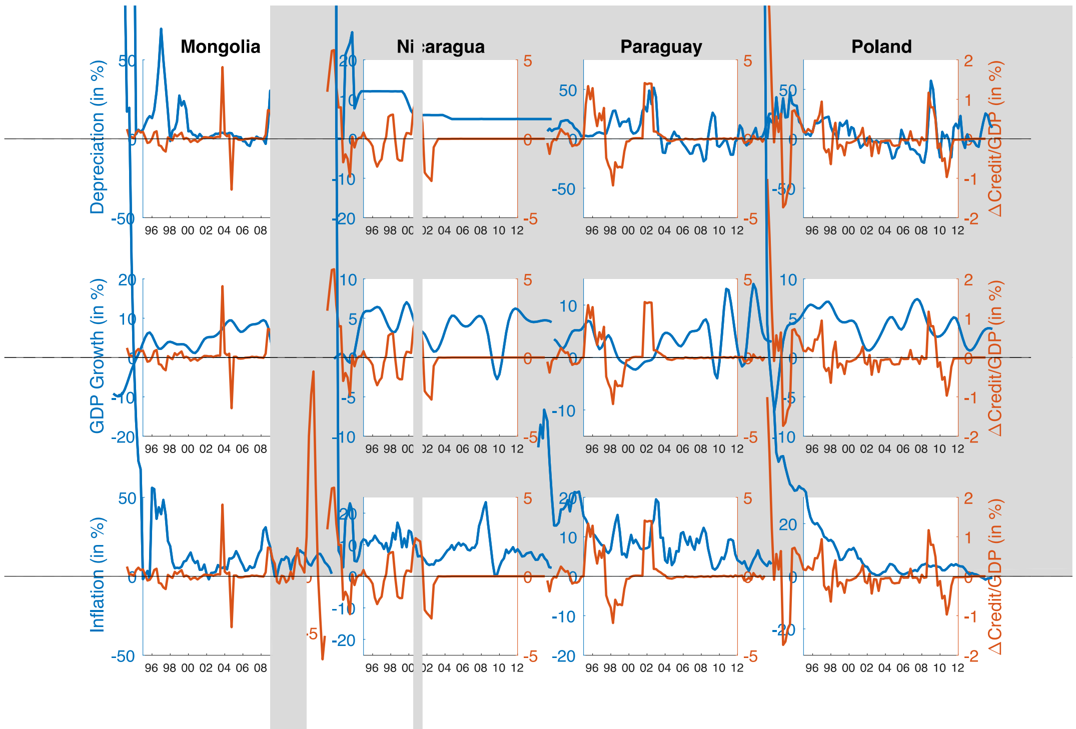
Shaded areas are banking crises as identified by Laeven and Valencia (2013).

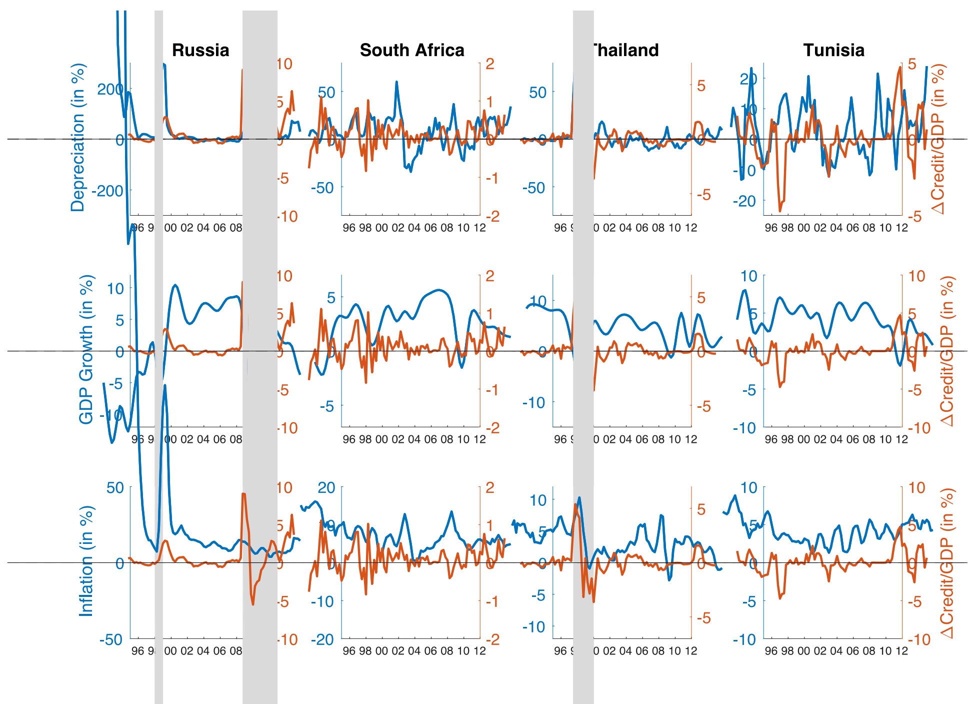


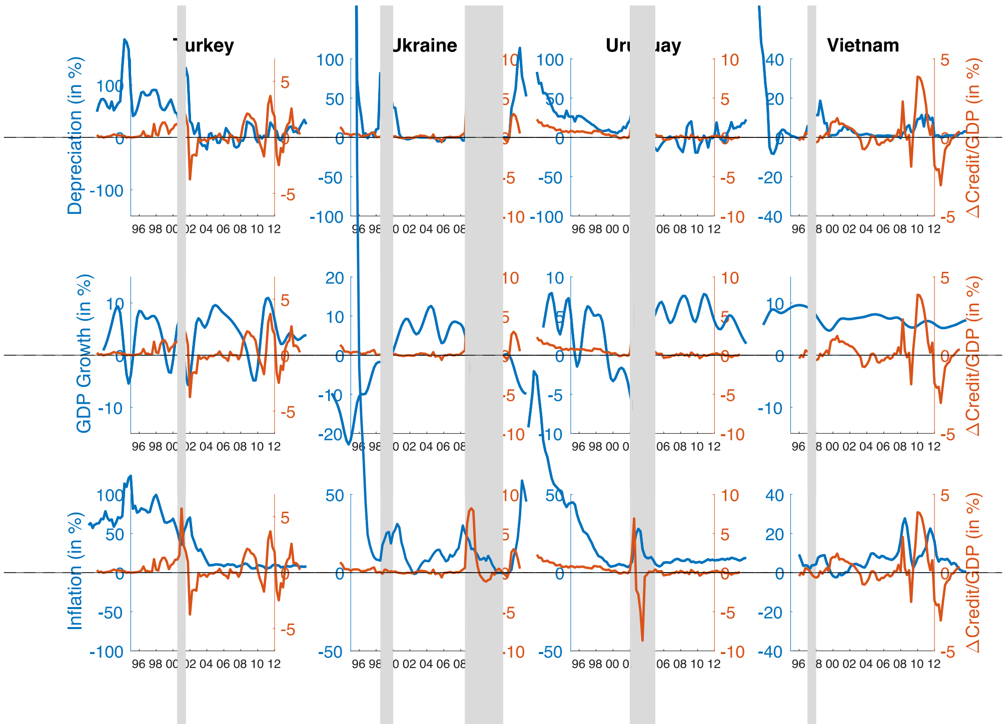


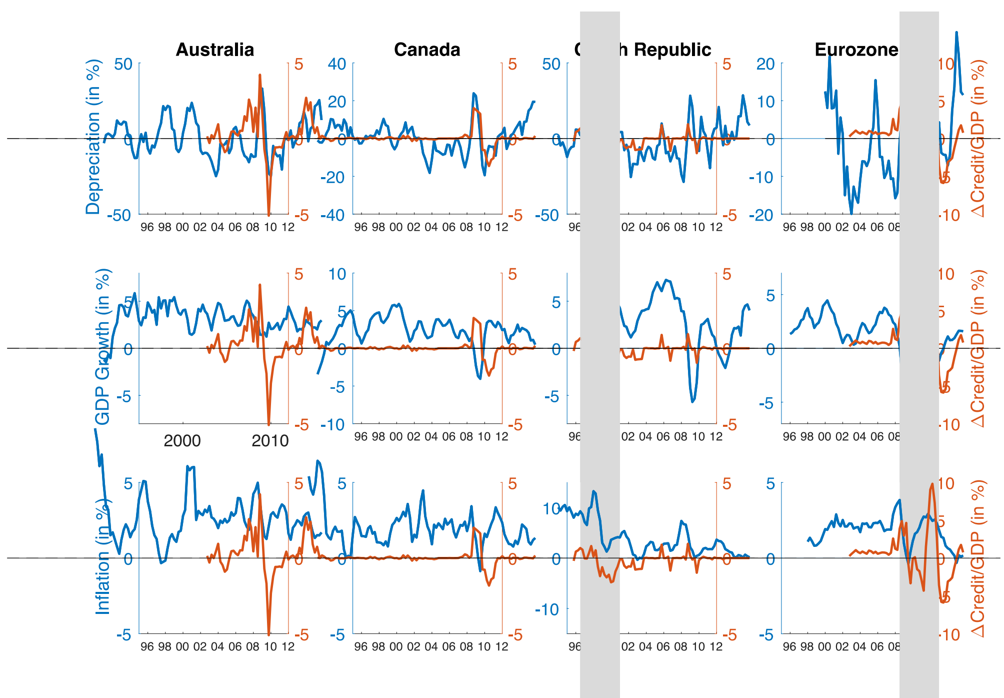


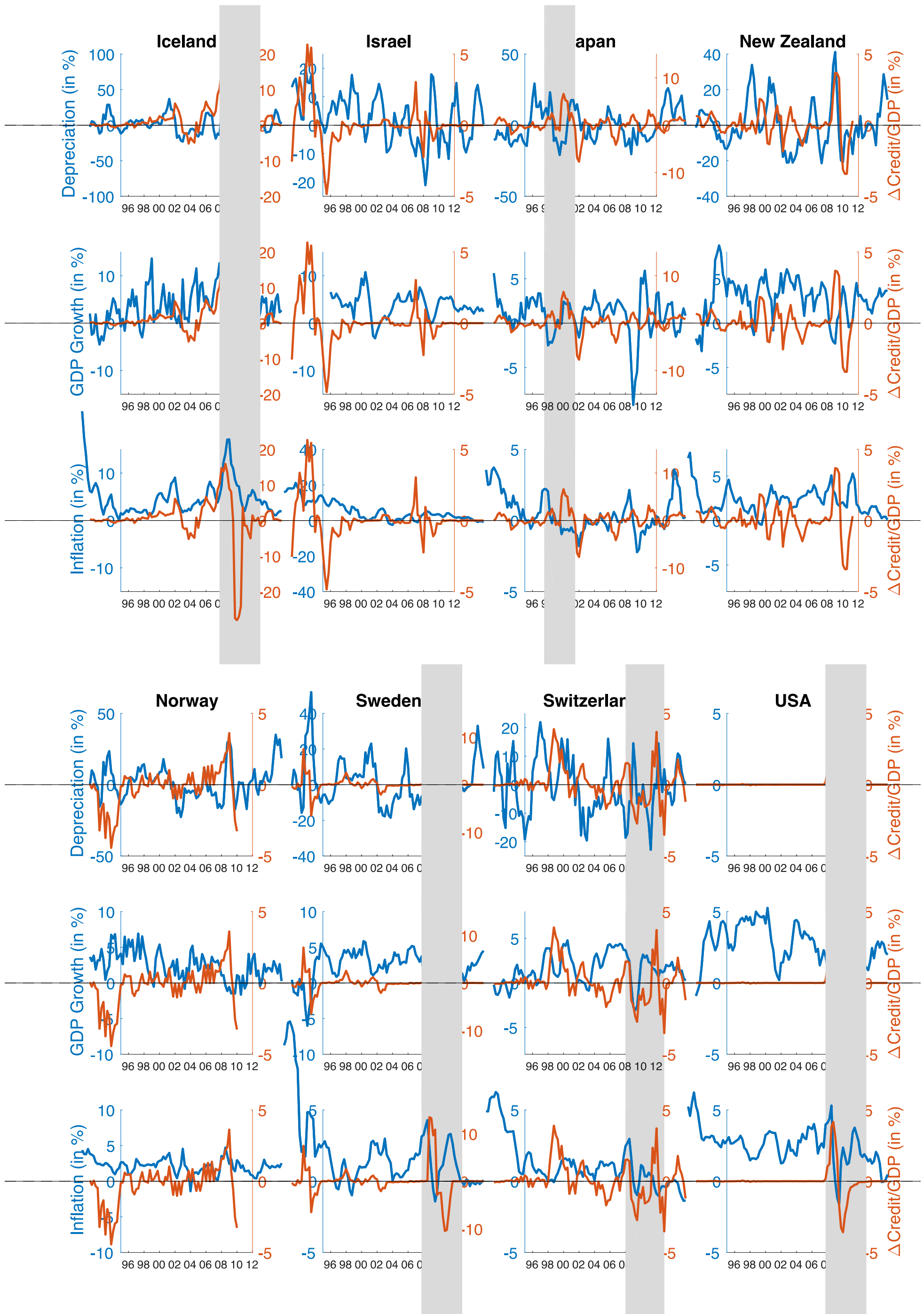
onesia
Kazakhstan
Kenya
Korea
yz Republic
Lithuania
Malaysia
Mexico
Depreciation (in %)
GDP Growth (in %)
Inflation (in %)
ΔCredit/GDP (in %)


Mongolia
Nicaragua
Paraguay
Poland
Depreciation (in %)
GDP Growth (in %)
Inflation (in %)
ΔCredit/GDP (in %)


Russia
South Africa
Thailand
Tunisia
Depreciation (in %)
GDP Growth (in %)
Inflation (in %)
ΔCredit/GDP (in %)


Vietnam
Depreciation (in %)
GDP Growth (in %)
Inflation (in %)
ΔCredit/GDP (in %)


Australia
Canada
Eurozone
Depreciation (in %)
GDP Growth (in %)
Inflation (in %)
ΔCredit/GDP (in %)



Sources: IFS, WEO.

Appendix V. Timing Restrictions in Local Projections

In this technical note, we show how to introduce short-run restrictions in the Local Projection (LP) framework. We provide proof of the claim that under short-run restrictions, structural impulse responses can be estimated one variable and one shock at a time.

As a starting point, let us define an impulse response (IR) as the difference between two forecasts (see, for instance, Hamilton 1994, and Koop and others 1996).

$$(A1)\; IR(h,d) = E(X_{t+h} \mid \epsilon_{k,t} = d) - E(X_{t+h} \mid \epsilon_{k,t} = 0), h = 0, 1, \dots, H$$

where E(.|) denotes the best, mean squared error predictor, X_t is a vector of N variables, $\epsilon_{k,t}$ is a *structural* shock, and $t = 0, 1, \dots T + H$.

Theoretically, when we can observe the true structural shock, we can directly obtain the structural impulse response of variable $X_{i,t}$ (the ith variable in X_t) to a structural shock $\epsilon_{k,t}$ from local projections, that is, we estimate

$$(A2)\; X_{i,t+h} = a^h + \beta^h \epsilon_{k,t} + u_{t+h}, h = 0, 1, \dots, H.$$

The structural impulse response functions are then given by $IR(h,d) = \beta^h d$. However, in practice we do not observe the true structural shocks and we need to postulate a structural model to identify them. For instance, assume that the economy can be described by the structural VAR representation of order 1.[22]

$$(A3)\; AX_t = F + BX_{t-1} + \epsilon_t,$$

As a simple normalization, let the diagonal elements of the matrix A all equal one. Now, imagine we are interested in the effects of a structural shock to variable $X_{k,t}$, that is, $\epsilon_{k,t}$. We can write the equation for variable $X_{k,t}$ as

$$(A4)\; A_k X_t = F_k + B_k X_{t-1} + \epsilon_{k,t},$$

where A_k, B_k and F_k are the k-th row of the A, B and F. Let us assume that there are M variables in X_t that react with a lag to $X_{k,t}$, and R variables that react contemporaneously. Therefore, N = M + R + 1. To simplify notation, let variables that react with a lag to $X_{k,t}$ be ordered before and variables that react contemporaneously after $X_{k,t}$. This implies that the last R entries of A_k are all equal to zero. We can solve for structural shock $\epsilon_{k,t}$:

$$(A5)\; \epsilon_{k,t} = A_k X_t - F_k - B_k X_{t-1},$$

or equivalently

[22] Note that this goes without loss of generality: we could postulate a VAR(p) instead and rewrite it in companion form. We use a VAR(1) for simplicity of exposition.

$$(A6)\ \epsilon_{k,t} = \sum_{j=1}^{N} A_{k,j}X_{j,t} - F_k - \sum_{j=1}^{N} B_{k,j}X_{j,t-1}$$

Using that the $A_{k,k} = 1$ and $A_{j,k} = 0\ \forall\ j > k$, we have

$$(A7)\ \epsilon_{k,t} = X_{k,t} + \sum_{j=1}^{M} A_{k,j}X_{j,t} - F_k - \sum_{j=1}^{N} B_{k,j}X_{j,t-1}$$

Plugging equation (A7) into (A2) yields the local projections

$$(A8)\ X_{i,t+h} = a^h + \beta^h \left[X_{k,t} + \sum_{j=1}^{M} A_{k,j}X_{j,t} - F_k - \sum_{j=1}^{N} B_{k,j}X_{j,t-1}\right] + u_{t+h}, h = 0, 1, \dots, H.$$

or rewritten for convenience

$$(A9)\ X_{i,t+h} = c^h + \beta^h X_{k,t} + \sum_{j=1}^{M} \gamma_j X_{j,t} + \sum_{j=1}^{N} \delta_j X_{j,t-1} + u_{t+h}, h = 0, 1, \dots, H.$$

By the virtue of our timing assumption we have that $E(X_{k,t}u_{t+h}) = 0$ and hence we can estimate the sequence of $H + 1$ projections in (A9) via OLS. Note that using $X = [\,Y, \pi, B, r, CE, E]$ and $X_k = CE$ yields expression (1) in the text using L = 1.

Timing Restrictions and State-Dependence in Local Projections

Consider now a state-dependent DGP:

$$(A10)\ (A + s_t A^s)X_t = F + (B + s_{t-1}B^s)X_{t-1} + \epsilon_t,$$

where the state s_t is the state-variable. Let us again use the normalization that the diagonal elements of A are equal to one. Due to the timing restriction, the last R entries in A_k and A_k^s are again equal to zero. Repeating the steps in (A4) – (A6) now yields

$$(A11)\ \epsilon_{k,t} = X_{k,t} + A_{k,k}^s s_t X_{k,t} + \sum_{j=1}^{M} A_{k,j}X_{j,t} + \sum_{j=1}^{M} s_t A_{k,j}^s X_{j,t} - F_k - \sum_{j=1}^{N} B_{k,j}X_{j,t-1} - \sum_{j=1}^{N} s_{t-1}B_{k,j}^s X_{j,t-1},$$

Plugging this into equation (A2), we get the local projection representation

$$(A12)\ X_{i,t+h} = a^h + \beta^h [X_{k,t} + A_{k,k}^s s_t X_{k,t} + \sum_{j=1}^{M} A_{k,j}X_{j,t} + \sum_{j=1}^{M} s_t A_{k,j}^s X_{j,t} - F_k$$

$$-\sum_{j=1}^{N} B_{k,j}X_{j,t-1} \ -\sum_{j=1}^{N} s_{t-1}B_{k,j}^{s}X_{j,t-1}] \ + \ u_{t+h}, h = 0, 1, \dots, H$$

We can simplify equation (12) to become

$$(A13)\ X_{i,t+h} = \ c^h + \ \beta^{h}X_{k,t} + \ \beta^{h,s}s_tX_{k,t} + \sum_{j=1}^{M} \gamma_j X_{j,t} \ + \ \sum_{j=1}^{M} s_t\gamma_j^{s}X_{j,t} - F_k$$

$$-\sum_{j=1}^{N} \delta_j X_{j,t-1} \ -\sum_{j=1}^{N} s_{t-1}\delta_j^{s}X_{j,t-1}] \ + \ u_{t+h}, h = 0, 1, \dots, H$$

By the virtue of our timing assumption we have that $E\left(X_{k,t}u_{t+h}\right) = 0$ and hence we can estimate the sequence of $H + 1$ projections in (A9) via OLS. Note that using $X = [\, Y, \pi, B, r, CE, E]$, $X_k = CE$ and s = B yields expression (2) in the text using L = 1.