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MONETARY POLICY, INFLATION, AND CRISES: NEW EVIDENCE FROM HISTORY AND ADMINISTRATIVE DATA

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Abstract

We show that U-shaped monetary policy rate dynamics are strongly associated with financial crisis risk. This finding holds both in long-run cross-country macro data covering many crises and monetary policy cycles, and in detailed micro, administrative data covering the post-1995 period in Spain. In the macro data, we find that pre-crisis monetary policy follows a U shape, with policy rates first cut and then increased over the 7 years before the onset of the crisis. This U shape holds across a wide variety of crisis definitions, short-term rate measures, and becomes stronger after World War 2. Differently, even though inflation and real rates show some of these dynamics before a crisis, these results are much less robust. The patterns are also much weaker when it comes to long-term rates and non-crisis recessions. We show that monetary policy rate hikes (both raw, and instrumented using the trilemma IV of Jorda et al, 2020) increase crisis risk, but, different to previous studies, we show that this effect is driven by rate hikes which were preceded by a series of cuts. To understand why U-shaped monetary policy is linked to crises, we show that the initial loosening of policy is followed by high growth in credit and asset prices, putting the economy into a vulnerable financial ``red zone''. After the subsequent monetary tightening these vulnerabilities materialize, leading to larger-than-usual declines in credit, asset prices, and real activity. To dig into the underlying mechanisms, we use administrative data on the universe of bank loans and defaults during the 1990s and 2000s boom-bust cycles in Spain. Consistently, we find that U-shaped monetary policy increases the probability of ex-post loan defaults, but effects are much stronger for ex-ante riskier firms and for banks with weaker balance sheets. Overall, our paper shows that monetary policy dynamics have important implications for financial stability.

JEL Classification: E51, E52, E44, G01, G21, G12

Keywords: monetary policy, Financial stability, Financial crises, Credit, Asset prices, Banks, Macro-finance

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Monetary policy, inflation, and crises: New evidence from history and administrative data ⋆

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1. Introduction

With annual inflation rates reaching 10% in 2022, central banks in Europe and the US started raising monetary policy rates to fulfill their price stability mandates. These policy actions, however, carry non-trivial trade-offs, with risks to macroeconomic growth and financial stability—especially after a period of low interest rates—featuring prominently in the surrounding debates (Acharya and Rajan, 2022, The Economist, 8th December 2022).

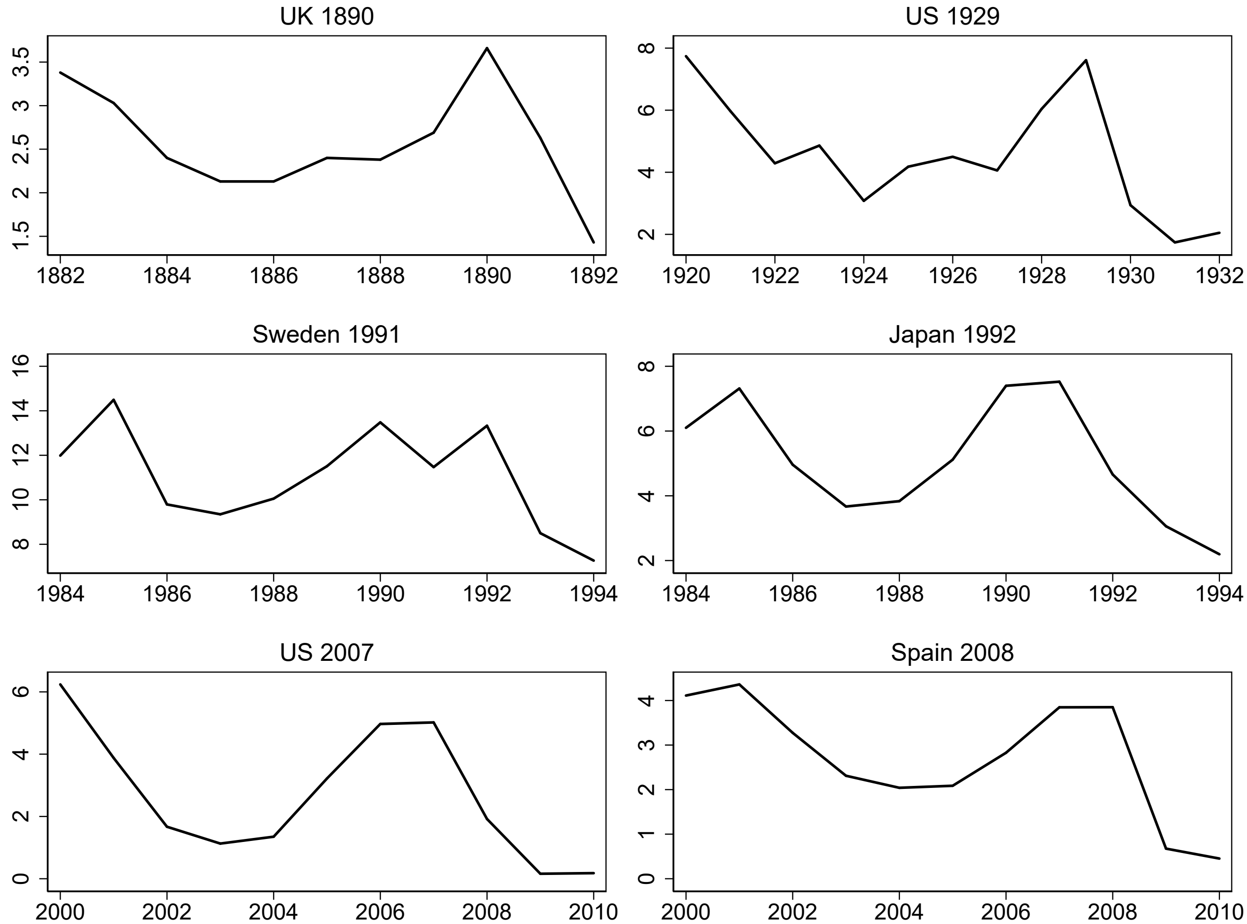
The trade-offs between monetary policy actions, inflation, and GDP growth are extensively researched and relatively well understood. But much less is known about how monetary policy and inflation affect financial stability, including the risk of a financial crisis. Existing studies on this topic offer relatively little in the way of consensus.[1] For monetary policy, there is evidence that low monetary policy rates foster risk taking by banks (e.g., Jiménez, Ongena, Peydró, and Saurina, 2014) and increase lending and asset prices (e.g., Kashyap and Stein, 2000; Bernanke and Kuttner, 2005; Hanson and Stein, 2015), two factors shown to be strongly associated with future crises (e.g., Greenwood, Hanson, Shleifer, and Sørensen, 2022). On the other hand, Schularick, ter Steege, and Ward (2021) find that it is policy rate hikes, rather than cuts, that increase crisis risk. For inflation, there is a long-standing argument that deflation increases the real value of debt and, with it, default risk (Fisher, 1933), but recent work by Agarwal and Baron (2021) also shows that unexpected increases in inflation negatively impact banking sector net worth. One reason why consensus in the literature is hard to reach may be that, when it comes to crisis risk, the whole path of past and current policy rates and inflation matters. For example, a period of loose monetary policy may foster excessive risky lending, and a subsequent policy tightening may trigger defaults on these risky loans.

In this paper, we study the link between the full path of monetary policy and financial stability. Experience of past crises suggests that the full path of monetary policy rates may indeed matter. Figure 1 shows the level of the monetary policy rate around important past crisis events: two famous global crises in the form of the 1930s Great Depression (shown for the US) and the 2007–08 Global Financial Crisis (shown for the US and Spain), and other more regional crises in the form of the UK Baring crisis in 1890, the Japanese crisis of the 1990s, and the 1990s Scandinavian crisis (shown for Sweden). In all these episodes, monetary policy displays one salient feature: the U-shaped pre-crisis policy rate path. In every one of these six crises, we observe sharp monetary policy rate cuts 8 to 3 years before the crisis, followed by rate increases in the years running up to the crisis onset.

In each of these episodes, the reasons for these policy rate cuts and increases were different. For example, in the US case studies, cuts in rates helped ensure a soft landing after the 1920–21 and 2000-01 recessions, and the subsequent increases were implemented to fight inflation, or stabilize the economy and the financial markets. In the cases of Japan, Spain, and Sweden, the policy rate path was mainly dictated by the need to maintain a fixed exchange rate peg or honor international

[1] For an extensive literature review on the theoretical mechanisms see Ajello, Boyarchenko, Gourio, and Tambalotti (2022), while the empirical evidence is summarized in Boyarchenko, Favara, and Schularick (2022).

Figure 1: *Monetary policy rates around selected past crisis events*



Notes: Level of the short-term nominal monetary policy rate for selected country and year. The crisis dates we use here follow Reinhart and Rogoff (2009b), while we will later employ the Jordà et al. (2016) chronology, which dates the banking crisis in the Great Depression to 1931 and the Japanese crisis to 1997.

agreements. Regardless of these underlying reasons, the narrative accounts of these financial crises often assign a key role to the path of monetary policy, with loose monetary policy contributing to the pre-crisis booms in credit and asset prices, and the subsequent tightening—to the reversal of these booms that eventually triggers the crisis (Taylor, 2013). Up to this point, however, this conjecture has not been systematically tested in the data.

To study the link between monetary policy cycles and financial stability, we exploit macro data for 17 countries going back to 1870 (from Jordà, Schularick, and Taylor, 2016), as well as detailed micro, administrative data covering the post-1995 period in Spain (from the Spanish credit register). This allows us to both cover many financial crises and monetary policy cycles in the macro analysis, and cover the universe of bank loans and defaults in micro data for one important financial crisis. In the cross-country analysis, we analyze the link between the full path of monetary policy and the probability of a subsequent banking crisis, as well as the effects on credit growth, asset prices and real variables (e.g., output growth and real estate investment).

We start by using the macro data to i). Characterize the full path of monetary policy and inflation around past crises, ii). Study the path-dependence in the links between policy rates and crisis risk, and iii). Assess the contribution of monetary policy to the well-documented boom-bust pattern in credit and asset prices around crises (Schularick and Taylor, 2012; Greenwood et al., 2022). We then turn to administrative data to zoom in on one of the episodes shown in Figure 1, Spain in the early 2000s, and analyze the heterogeneous effects of the path of monetary policy on individual loan defaults across firms and banks. To do this, we test whether loans granted during periods of loose monetary policy are more likely to default when policy is tightened, and if these effects are stronger for specific economic sectors (e.g., real estate), and for riskier banks and firms.

We first document the full path of monetary policy rates, inflation, and real rates around different financial crisis episodes. We find that the patterns documented for the six case studies in Figure 1 extend to the average crisis in the long-run sample. Short-term policy rates decrease from about 7 years prior to a crisis, before they start rising in the 3 years running up to the crisis. This pattern holds across crisis definitions (Jordà et al., 2016 and Baron, Verner, and Xiong, 2021), for the subset of deep crises (i.e., those accompanied by low GDP growth), and in the post-1945 sample. Binning the data into 2×2 different policy shapes, we find that the frequency of financial crises over a three-year window is 20% following an eight-year period characterized by a policy rate cut over the first five years, and a subsequent increase over the following three years. This is more than twice the frequency after other policy shapes (double raises, double cuts, a raise and cut Λ), and twice the unconditional crisis probability in a 3-year window.

These differences become even more pronounced when we look at the subsets of deep crises and post-WW2 crises only. In fact, for our sample of advanced economies, every single deep crisis after World War 2 was preceded by a U-shape in interest rates. When we look at inflation and real rates, all these patterns are less clear, with large heterogeneity across different sample periods. While real rates show some U pattern in the pre-1945 sample, this relationship disappears with the advent of systematic monetary policy used to stabilize the macroeconomy after 1945. It is important to highlight that there are no expected inflation data covering our historical sample of crises, so we focus on ex post rather than ex ante real rates. Further, effects are driven by short-term (monetary policy) rates rather than long-term interest rates.

Another way to think about this U shape before crises is that policy rate hikes entail different crisis risks depending on the previous path of monetary policy rates. When we put this logic into a crisis prediction framework, we find that monetary policy rate increases are associated with a higher likelihood of a subsequent crisis (as in Schularick et al., 2021), but that this relationship is almost entirely due to episodes when interest rates were previously cut. To mitigate potential endogeneity concerns in such regressions, we use the Jordà, Schularick, and Taylor (2020) trilemma instrument. This instrumental variable approach exploits that with open capital accounts and fixed exchange rates countries have to give up monetary autonomy—the famous trilemma of international finance.

The IV results confirm our hypothesis that the effect of monetary policy on crisis risk is highly path dependent. Three-year increases in monetary rates only predict crises if rates had been cut over the five-year window before. The effects that we find are both economically and statistically significant: a cut in rates during years $t-8$ to $t-3$ followed by a 1 percentage point increase in rates during years $t-3$ to t increases the probability of a crisis between years t and $t+2$ by 13 percentage points, more than doubling crisis risk. These effects are also much larger than the 3 percentage point increase in crisis probability after an unconditional 1 percentage point increase in monetary policy rates (i.e., not conditioning on a previous cut).

We also analyze the mechanisms for our main result. As the narrative accounts for our case studies in Figure 1 suggest, the reason for this strong state-dependency may lie in the boom in financial markets that is triggered by the previous cut in rates. When we study the behavior of asset prices and credit after monetary policy rate changes, we find evidence consistent with this channel. A cut in policy rates is followed by credit booms and rising valuations of housing and equity. As a combined measure of these vulnerabilities, we follow Greenwood et al. (2022) and define an economy to be in the financial red zone ("R-zone") if credit and asset prices are jointly elevated. It turns out that a cut in policy rates makes the economy more likely to end up in the R-zone.

Moreover, adding the second leg of the U, we then study the adjustments when monetary policy rates increase in this environment. We find that the drops in credit and asset prices are much stronger if policy rate hikes were preceded by a cut. For example, an instrumented 1 percentage point increase in monetary policy rates leads to 1 percentage point lower cumulative real credit growth over the next 3 years if rates were previously raised, and 3 percentage points lower credit growth if rates were previously cut. Similarly, the likelihood of a financial crisis increases when monetary policy rates increase, and the economy is in the R-zone. We find similar state dependencies for measures of real activity: increases in rates that were preceded by cuts lead to much larger declines in real GDP, consumption, and housing investment.

We then turn to the Spanish credit register data to dig deeper into the underlying mechanisms. In particular, we exploit information on ex-post default at the loan level to study, again, the effects of U-shaped monetary policy rate dynamics on financial risk. In addition to having the universe of bank loans and defaults, the advantage of using Spanish credit register data in this context is that the behavior of policy rates closely mirrors our instrumental variable strategy in the aggregate data. With Spain joining the European Monetary Mechanism in 1989 and being part of the Eurosystem after 1999, monetary policy rates were decided in Frankfurt, not in Madrid. Before the Euro Area Sovereign Debt Crisis, Spanish monetary policy therefore followed the euro area core countries with different economic dynamics from Spain. This allows us to study the implications of monetary policy rates for loan defaults when policy rates display a U shape, including heterogeneities across firms and banks.

We find that a monetary policy rate cut over the previous 5 years ($t-5$ to t) is associated with a significantly higher likelihood of loan default at time $t+1$. At the same time, contemporaneous (at time $t+1$) increases in monetary policy rates are also associated with a higher likelihood of loan default. Interacting the two, monetary policy rate hikes following a cut are associated with a significantly higher likelihood of default. A rate cut over a 5-year period followed by a 1 percentage point raise in rates increases loan default probability by 28.4% in relative terms. This is roughly double the unconditional increase in default probability after a 1 ppt monetary policy rate increase. These results hold across a wide range of specifications, including several layers of fixed effects to control for bank, borrower, and time characteristics. We also find that these patterns are more pronounced for riskier firms, short-term loans, loans extended by weak banks, and loans extended to the construction and real estate sector.

The paper cuts across several strands of the literature, linking the seemingly ambiguous results on the relationship between monetary policy rates and financial risk as well as combining evidence from aggregate long-run and granular loan-level data. The closest strand of the literature is the macroeconomic evidence on the link between monetary policy rates and financial stability. Here, there is seemingly conflicting evidence that interest rate cuts, on the one hand, increase credit growth and asset prices—two leading sources of financial instability pointed out by previous research (Bernanke and Kuttner, 2005; Maddaloni and Peydro, 2011; Brunnermeier and Sannikov, 2014; Gertler and Karadi, 2015; Hanson and Stein, 2015; Dell'Ariccia, Laeven, and Suarez, 2017; Adrian, Duarte, Liang, and Zabczyk, 2020)—but on the other, when they study the link between monetary policy rates and crises directly, Schularick et al. (2021) show that it is monetary rate increases, not cuts, that increase crisis risk. Existing theoretical models also point to potential links between low interest rates, asset prices, and credit (e.g., Stein, 2012) but there is little research on the theoretical links between the full path of monetary policy and crisis risk. As an exception, a recent paper by Boissay, Collard, Galí, and Manea (2021) shows that a sequence of low interest rates followed by surprise rate hikes leads to a higher crisis risk, in line with our empirical results.

We also contribute to the literature on monetary policy, credit provision, and risk taking. This literature tends to find that monetary policy easing increases lending (e.g., Kashyap and Stein, 2000; Jiménez, Ongena, Peydró, and Saurina, 2012; Drechsler, Drechsel, Marquez-Ibanez, and Schnabl, 2016; Acharya, Imbierowicz, Steffen, and Teichmann, 2020) and encourages bank risk-taking (e.g., Adrian and Shin, 2010; Jiménez et al., 2014; Martinez-Miera and Repullo, 2017). Similarly, there is evidence of reach for yield in other financial intermediaries (e.g., Becker and Ivashina, 2015; Di Maggio and Kacperczyk, 2017). We contribute to this literature by showing that the full path of monetary policy rates (in particular, the U-shaped rate path) affects financial stability, and going beyond the micro-level analysis of loan defaults and reach for yield to also study the probability of financial crises, as well as credit and asset price dynamics at macro level.

A large literature has studied the link between monetary policy and the macroeconomy (see Ramey, 2016 for an overview), and several papers have argued in favor of important state dependencies (e.g., Tenreyro and Thwaites, 2016; Jordà et al., 2020). Berger, Milbradt, Tourre, and Vavra (2021) show that one such state dependency may actually lie in the past path of monetary policy itself, arguing that recent monetary policy decisions in the US may have affected the sensitivity of real activity to future rate changes through mortgage prepayment and, in effect, might have limited future policy space. A similar channel is explored in Eichenbaum, Rebelo, and Wong (2022).

While these papers think about this path dependence mostly in terms of the ability to stimulate the economy through interest rate cuts, our results suggest that a similar state-dependency might apply for the sensitivity of financial stability to future rate hikes.

Finally, our results contribute to the large literature on financial crises. This literature has shown that financial crises entail high macroeconomic costs (e.g., Cerra and Saxena, 2008; Reinhart and Rogoff, 2009a). At the same time both crises and their severity are predictable using "early warning indicators"—such as credit expansions and asset price growth (e.g., Schularick and Taylor, 2012, Borio and Lowe, 2002, Greenwood et al., 2022)—and measures of financial vulnerabilities—such as elevated asset prices, leverage, and bank equity losses (e.g., Jordà, Schularick, and Taylor, 2013; Jordà, Schularick, and Taylor, 2015; Jordà, Richter, Schularick, and Taylor, 2021; Baron, Verner, and Xiong, 2021). Our paper contributes to this literature in identifying low monetary policy rates as one of the sources for the buildup of these vulnerabilities, and by adding the reversal in interest rates as a trigger for the subsequent loan defaults and crisis dynamics.

The remainder of the paper is organized as follows. We document the full path of monetary policy rates and inflation around crises in Section 2. We analyze the relationship between the dynamic shape of monetary policy and crisis risk in Section 3. We study the underlying mechanisms in macro data in Section 4, and drill into the heterogeneities across firms and banks using Spanish credit register micro data in Section 5. Section 6 concludes with some policy implications.

2. Monetary policy and inflation around crises

2.1. Main results

The case studies presented in the introduction display a very pronounced U shape for monetary policy prior to financial crises. Without further analysis it is, however, difficult to tell, first, whether this pattern of monetary policy is systematic, i.e., indicative of monetary policy around crises in general, and second, whether there is a specific mechanism linking U-shaped monetary policy to crisis risk. In the remainder of the paper, we will explore both of these questions, starting with the empirical analysis of monetary policy shape around past crisis events in Sections 2 and 3, and moving on to study the underlying mechanisms in Sections 4 and 5. To establish the stylized patterns regarding the shape of policy and crises, we will rely on historical data in the Jordà et al. (2016) macro-history database, which covers 17 advanced economies between 1870 and 2020, with around 90 systemic banking crisis events (72 crises for the sample that includes the full path of nominal rates and inflation around an extended crisis window). To dig into the mechanisms in Sections 4 and 5, we will use a mix of these macroeconomic data covering the representative sample of crises, and administrative-level credit registry data for Spain covering the 1990s and 2000s macro-financial boom-bust cycles in this country.

Crisis-window averages We start by plotting the average levels of nominal policy rates, inflation, and real interest rates around past crises using simple event windows. To do this, we simply take the level of each of these variables at $t-7$ to $t+7$ around each crisis date t, and average these levels across the 72 crises in our sample. The results are shown in Figure 2. The solid black lines show the average level of the short-term interest rate (policy rate), inflation, and real interest rate (short-term rate minus inflation in the same year), for the window from 7 years before to 7 years after the onset of the average crisis. The dashed green lines show the average level of these variables across all observations that do not coincide with the crisis start date. Panel (a) uses the Jordà et al. (2016) crisis chronology, which catalogues systemic banking crises using a narrative panic definition. Panel (b) compares these results to (i) the alternative chronology of Baron et al. (2021) (BVX crises), who use both narrative panics and large declines in bank stock returns to classify crises, and to (ii) a "deep crisis" definition which considers only those crises that are accompanied by low GDP growth (-3% in one year, or -1% on average over 3 years, in the $t-1$ to $t+3$ window around the crisis start date). Panel (c) shows the averages for each of these types of crises during the post-1945 time period.

The left panel of Figure 2a shows that the average crisis is preceded by U-shaped monetary policy, with a series of cuts in interest rates 7 to 4 years before the crisis followed by increases during the three years in the run-up to the crisis. Panels (b) and (c) show that the U shape also holds for the Baron et al. (2021) crisis definition, and is more pronounced for deep crises, especially after World War 2. The swings in policy rates are substantial, around 1 ppt down and up over the full sample, and as large as 3 ppts on average for deep crises starting after World War 2. The middle and right panels of Figure 2 show the average levels of inflation and real interest rates around crises. There is some evidence that crises are preceded by spikes in inflation and a U shape in real interest rates (panels (a) and (b)), but these patterns become much weaker or reverse after World War 2 (panel (c)). This suggests that the U shape in nominal interest rates is an overall more robust feature of crises than the inflation spike, and the U shape in real rates.

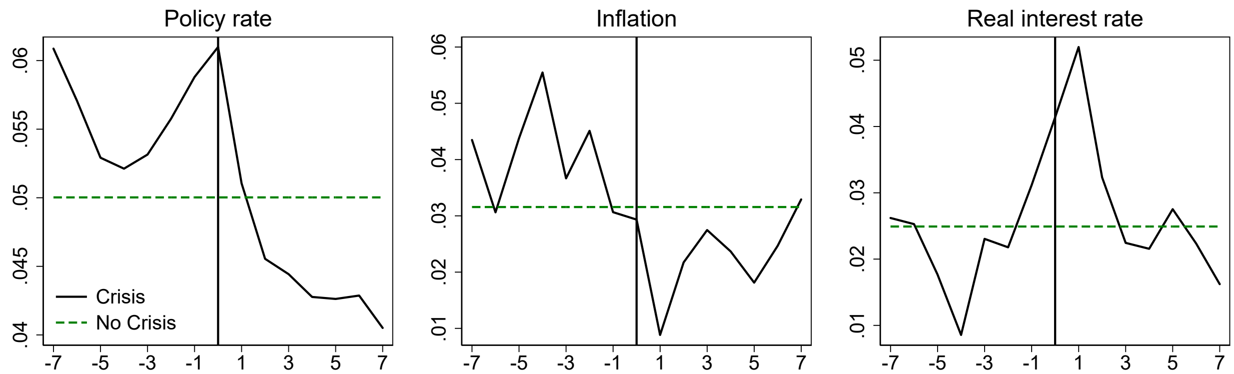
Crisis window regressions To test whether these pre-crisis patterns of interest rates and inflation are statistically significant, we run the following regression:

$$y_{i,t+h} - y_{i,t} = \alpha_{i,h} + \alpha_{d,h} + \beta_h \mathbb{1}_{Crisis_{i,t}=1} + \epsilon_{i,t+h}. \tag{1}$$

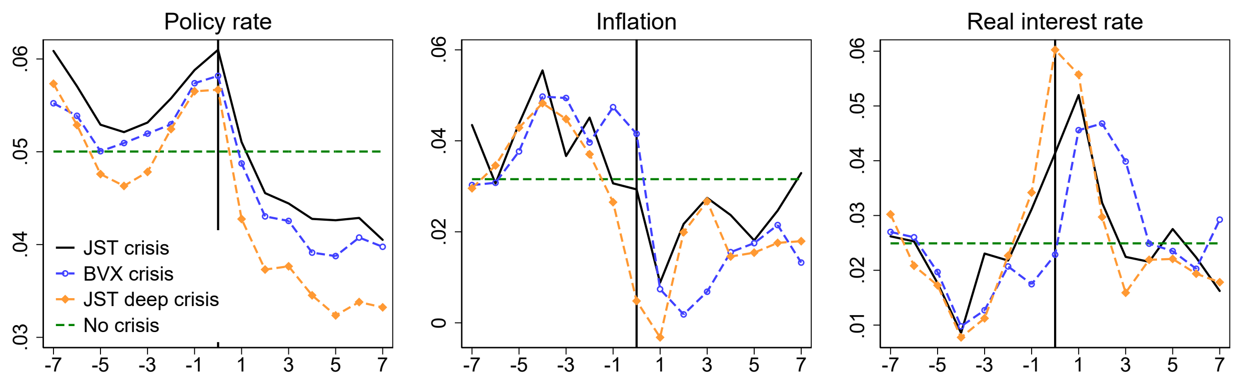
Above, y are policy rates, inflation, and real rates, $h = -7, ..., 0, ...7$ are horizons with $y_{i,t+h} - y_{i,t}$ corresponding to changes between 7 years before year t ($h = -7$) and year t, 7 years after year t ($h = 7$) and year t, and so on. $\mathbb{1}_{Crisis_{i,t}=1}$ is an indicator for the start of a systemic banking crisis in country i in year t, and α_i and α_d are country and decade fixed effects. We add decade fixed effects to control for common cross-country trends in the observed variables, such as the decline in interest rates after 1980 (Holston, Laubach, and Williams, 2017) and the "great moderation" of the 1990s (Stock and Watson, 2002). Figure 3 shows the estimated β_h coefficients for different horizons h alongside the 90% confidence intervals, for different outcome variables, crisis definitions, and

Figure 2: *Monetary policy rates and inflation around past crises*

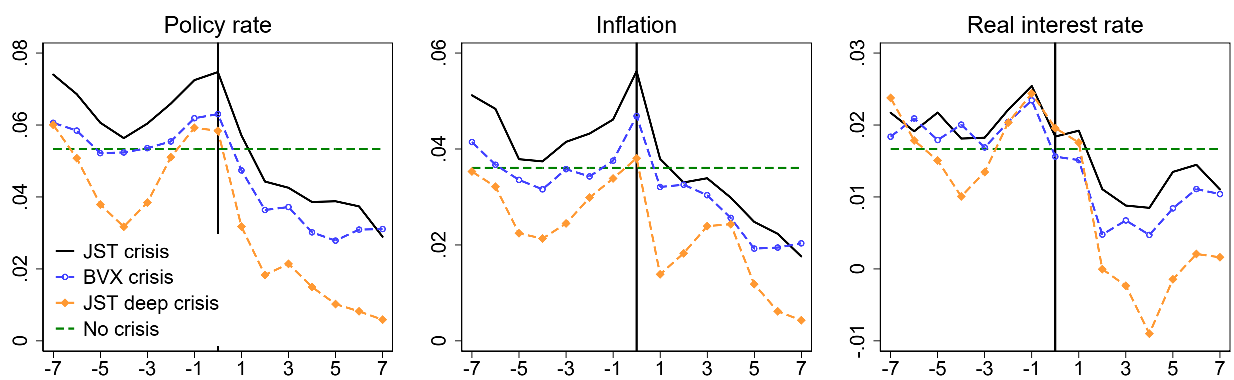
(a) *Jordà et al. (2016) systemic banking crises*



(b) *Alternative crisis definitions*



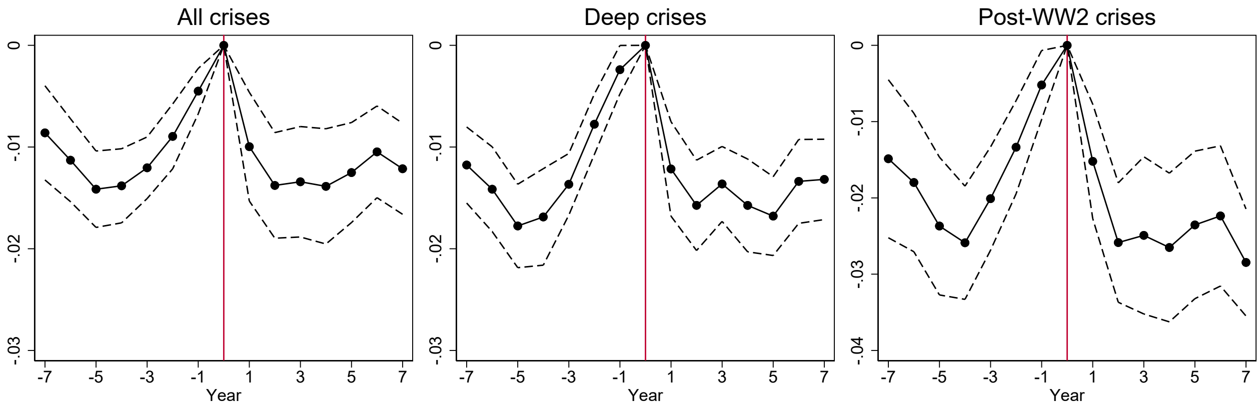
(c) *Post World War 2 crises*



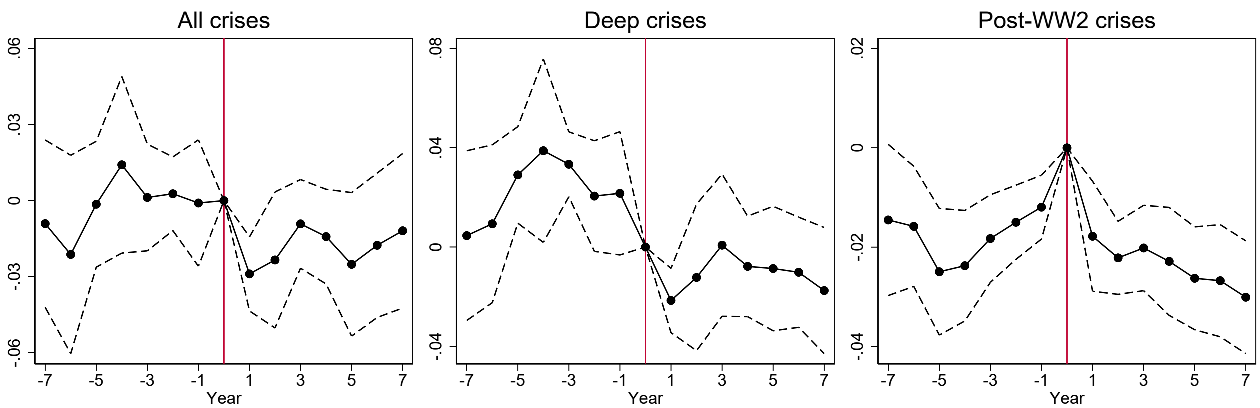
Notes: Unweighted averages of the level of the corresponding variable in year t (start of the crisis at $t = 0$). Total of 72 crises (24 post-WW2). Panel (a) uses the narrative crisis definition from Jordà et al. (2016). Panel (b) additionally considers the Baron et al. (2021) crisis chronology (BVX crises) based on bank equity returns, and deep crises (JST deep crises) defined as Jordà et al. (2016) banking crises with -3% or less GDP growth in one year, or average -1% or less GDP growth over 3 years in the $t - 1$ to $t + 3$ crisis window. Panel (c) limits the sample to crises that started after 1945. Green dashed lines show the mean of the respective variable for non-crisis observations.

Figure 3: *Policy rates and inflation: Crisis window regressions*

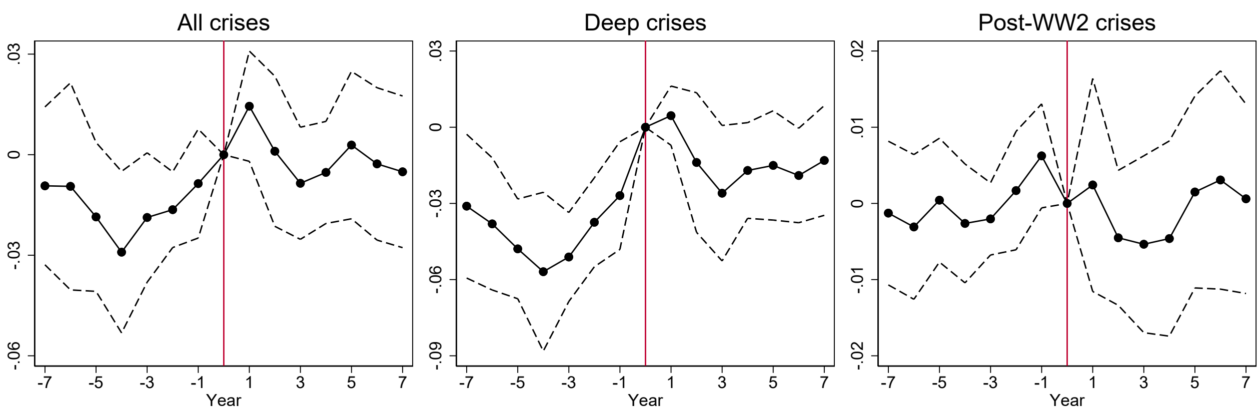
(a) *Monetary policy rates*



(b) *Inflation*



(c) *Real interest rates*



Notes: These graphs show the regression coefficients and 90% confidence intervals from regressing policy rates (panel a), inflation (panel b), and real rates (panel c) on the crisis dummy for horizons $h = -7, ..., 0, ...7$, with 0 corresponding to the beginning of the crisis according to the Jordà et al. (2016) chronology. Deep crises are those with -3% or less GDP growth in one year, or average -1% or less GDP growth over 3 years in the $t-1$ to $t+3$ crisis window. Post-WW2 crises are those that started after 1945. Inflation is the change in the CPI index. Real rate is the difference between the nominal rate and inflation in the same year.

samples. On average, all the variables follow the same patterns as in Figure 2. Importantly, the path of nominal policy rates is measured with much more precision than that of inflation and real rates, which again suggests that the U shape in nominal rates is a much more prominent precursor of crisis than the spike in inflation and the U shape in real rates.

2.2. Robustness

Residual rates Does the U-shaped nominal rate path simply reflect monetary policy responses to other developments in the economy? To look at this, we separate policy responses into endogenous and residual by running different versions of the Taylor rule, and plot their path around crises using the window regression in (1). Appendix Table A.1 shows the estimated Taylor rules, and Appendix Figure A.1 shows how the Taylor rule residuals behave around crises. In addition to the residuals, we also plot the cumulated changes in the Jordà et al. (2020) trilemma instrument—the changes in interest rates to defend the exchange rate peg—in the Appendix Figure A.1c. All of these different interest rate measures follow a U shape. This suggests that the path of monetary policy before crises goes beyond mechanical responses to observed macro-financial variables.

Long-term rates Does the U shape policy path also manifest itself in the long-term rates? To check this, Appendix Figure A.2a plots the crisis window regression coefficient for the long-term government bond yield (maturity of about 10 years), using the same scale as that for the short-term rate in Figure 3a. The response of the long-term rate is much weaker than that of the short-term rate. Appendix Figure A.2b shows the path of the term premium, the difference between the long and short-term rates. Before crises, the term premium actually follows a Λ-like pattern, the reverse of the U, with the gap between long and short-term rates increasing from $t-7$ to $t-4$ and shrinking from $t-3$ to t. This again confirms that the response of long-term rates is relatively more muted compared to that of short-term rates.

Recessions Appendix Figure A.3 shows the levels of policy rates, inflation, and real rates around non-financial recessions, obtained by running recession window regressions. To do this, we use the same set-up as for crises in equation (1), but use the business cycle peak as the dummy on the right-hand side (instead of the crisis dummy). It turns out that the patterns documented in Figures 2 and 3 are mostly specific to crises. Inflation and real rates show no clear patterns around non-financial recessions. Nominal monetary policy rates increase in the years before an average non-crisis recession, but the magnitude of the increase is much smaller than for crises (less than 10 basis points), and there is little evidence of a decline in years $t-7$ to $t-3$ before the non-crisis recession.

Table 1: *Monetary policy shape and crisis frequencies*

	(1)	(2)	(3)	(4)
	Crisis	Deep crisis	Post-WW2 crisis	Post-WW2 deep crisis
U shape (cut, raise)	0.20	0.13	0.18	0.14
Raise, raise	0.08	0.04	0.03	0.00
Raise, cut	0.05	0.02	0.01	0.00
Cut, cut	0.04	0.02	0.02	0.00
Unconditional	0.10	0.05	0.06	0.03

Notes: This table reports the crisis probability between year t and $t+2$ for different crisis definitions and paths of nominal monetary policy rates. Crisis frequency is the ratio of crisis to total observations in those years. Crises are dated using the Jordà et al. (2016) chronology. Deep crisis is crisis accompanied by at least -3% GDP growth in 1 year, or -1% average growth over 3 years in the window $t-1$ to $t+3$ around the crisis. Post-WW2 crises are those which started after 1945. In rows, the 4 bins are defined by the sign of the change (cut or raise) in the nominal policy rate between $t-8$ and $t-3$ and the sign of the change (cut or raise) between $t-3$ and t. For example U-shape (cut, raise) refers to a cut in rates between $t-8$ and $t-3$ and a subsequent raise between $t-3$ and t.

3. The path of monetary policy and crisis risk

What is the link between the dynamic path of monetary policy and crisis risk? We study this question in two steps. First, we compare crisis frequencies after different monetary policy paths, to see which type of policy path is associated with a higher risk of crises. Second, we run crisis prediction exercises to test if changes in monetary policy rates predict crises, and if this predictive power changes depending on the past path of monetary policy.

3.1. Policy rate paths and crisis frequencies

Nominal policy rates We start with a simple analysis of crisis frequencies after different paths of monetary policy. Figure 2 showed that on average, crises are preceded by U-shaped monetary policy, with several years of cuts in rates followed by several years of increases. But it could still be the case that some other policy paths result in similarly high crisis frequencies compared to the U shape. To check if this is the case, we test whether crises are more frequent after U-shaped monetary policy than other policy paths. To do this, we classify pre-crisis monetary policy into four paths, based on the monetary policy rate changes in the 8 years before the crisis. We divide these 8 years into a 5-year period that is supposed to capture an extended period of decreasing policy rates, and a subsequent 3-year window that allows us to capture the reversal in interest rates.[2] First, we classify as the U shape any policy actions that resulted in a cumulative rate cut between years $t-8$ and $t-3$, followed by a raise from $t-3$ to t. Second, we consider double raise episodes of subsequent increases from $t-8$ to $t-3$, and from $t-3$ to t. Finally, we have the Λ shape of raises from $t-8$ to

[2]Constant rates are classified as increases, but results do not change if we classify them as cuts.

$t-3$ and cuts from $t-3$ to t, and double cut from $t-8$ to $t-3$, and $t-3$ to t.

We then compute the post-path crisis frequency by taking the ratio of crisis observations to all observations with non-missing data after the end of each specific monetary policy rate path. Since the exact timing of the onset of a crisis is difficult to predict, as our baseline definition we consider all crises in the three-year window from t to $t+2$ after the end of the policy shape at t.[3] The corresponding crisis frequencies, for different crisis definitions, are shown in Table 1. We consider the baseline Jordà et al. (2016) crisis definition in columns 1 (full sample) and 3 (post-WW2), and the deep crises which are accompanied by low GDP growth in columns 2 and 4.

U-shaped monetary policy is associated with a substantially higher risk of systemic banking crises. If rate increases unambiguously increased crisis risk, we would expect the double raise in row 2 of Table 1 to be associated with the highest risk of crises, but this is not the case. In the data, crises are more than twice as likely after the U shape than after a double raise (20% versus 8% crisis probability in column 1), and twice the unconditional probability of experiencing a banking crisis during these 3 years (10%). These differences become even larger when we consider deep crises in Table 1 column 2, and post World War 2 crises in column 3. When it comes to deep systemic banking crises which started after World War 2 (column 4), every single one of these was preceded by a U shape in nominal monetary policy rates, and none by other policy shapes.

Appendix Table A.2 shows that these results hold if we consider a narrower 1-year crisis window instead of the 3-year window shown in Table 1. Appendix Table A.3 shows that the monetary policy rate U shape is not as prominent a feature of non-crisis recessions, compared to financial crises: for example, after World War 2 a deep non-financial recession is more likely after a double raise than after a U shape in monetary policy.

Inflation and real rates The association between nominal monetary policy rates and crisis frequencies is highly path dependent. Is the same true for inflation and real interest rates? Appendix Table A.4 plots the crisis frequencies for different paths of real interest rates, and Appendix Table A.5 does the same for inflation. As with nominal rates, we classify a U shape in real rates if the real interest rate (or inflation) is falling between $t-8$ and $t-3$ and increasing between $t-3$ and t, and so on. There is little difference between crisis probabilities for different paths of inflation. There are some differences in crisis frequencies across real interest rate paths. In particular, deep crises are much more frequent after the U-shaped real rate path than after other paths of the real interest rate. But overall these differences are much less stark than those for nominal rates.

[3] A few of the crises can be associated with different policy shapes depending on where we are in the 3-year crisis window. In these cases, we assign the crisis to the most frequent of these policy shapes within the window. For example, if the crisis is associated with a U in the 8 years preceding time t and $t+1$, and a double raise in the 8 years preceding $t+2$, we classify it as being preceded by a U.

3.2. Predicting financial crises

Nominal policy rates We next evaluate whether the U shape also allows to predict crises. To do this, we regress a crisis dummy on the change in the monetary policy rates, and allow the effects of changes in rates on crisis risk to vary depending on the past path of monetary policy by additionally including a dummy for whether rates were cut in the past, and the interaction of the change in rates with this dummy. To be consistent with results on pre-crisis window averages in Section 2, and the crisis frequencies in Section 3.1, we use the window of $t-8$ to $t-3$ to define a dummy for the cut in monetary policy rates, the window of $t-3$ to t for the rate increase, and the window of t to $t+2$ for the dummy indicating the start of the systemic banking crisis. Specifically, we estimate linear probability models for a crisis occurring between t and $t+2$ as follows:[4]

$$\text{Crisis}_{i,t \text{ to } t+2} = \alpha_i + \beta_1 \Delta_3 \text{Rate}_{i,t} + \beta_2 \text{Cut}_{i,t-8,t-3} + \beta_3 \Delta_3 \text{Rate}_{i,t} \times \text{Cut}_{i,t-8,t-3} + \gamma X_{i,t} + u_{i,t \text{ to } t+2}. \quad (2)$$

Above, i and t are country and year indices, Δ_3Rate refers to three-year changes in monetary policy rates, and Cut takes the value of 1 if rates were cut (cumulative change < 0) between years $t-8$ and $t-3$. $X_{i,t}$ is a control vector that includes eight lags of real GDP growth and inflation.

We run the above regressions both for raw changes in rates Δ_3Rate, and for changes in rates instrumented using the trilemma IV of Jordà et al. (2020). The intuition behind the trilemma IV is that countries in fixed exchange rate regimes with open capital accounts are forced to track monetary policy in the base country (Mundell, 1963). This means that base country interest rate changes can be used as instruments for changes in interest rates in the peg country. To this end, in the IV specifications, we follow Jordà et al. (2020) and instrument the change in interest rates in a country in a fixed exchange rate regime with the change in rates in the base country residualized conditional on economic conditions in the base country, as follows:

$$\text{Trilemma IV}_{i,t} = (\Delta \text{Rate}_{b(i),t} - \widehat{\Delta \text{Rate}}_{b(i),t}) * PEG_{i,t} * PEG_{i,t-1} * KOPEN_{i,t}. \quad (3)$$

Above, ΔRate_b is the change in nominal policy rates in the base country (e.g., Germany for countries in the ERM mechanism), residualized by subtracting the value of $\widehat{\Delta \text{Rate}}_b$ predicted by the economic conditions in the base country (inflation, GDP, credit, asset prices, consumption, investment, lagged interest rates), PEG variables are dummies for fixed exchange rate regimes, and KOPEN measures the degree of capital account openness. To mirror the set-up of equation (2), we instrument the three-year changes in rates and their interaction with the cut dummy with, respectively, the three-year change in the residualized trilemma instrument and its interaction with the cut dummy.

The results of these regressions are shown in Table 2. Column 1 shows the results of regressing the crisis dummy on three-year changes in monetary policy rates in a simple OLS regression with

[4]Note that we allow here for a contemporaneous relationship between raises in interest rates and financial crisis as indicated by our case studies. In robustness exercises in the appendix we confirm that these relationships also hold when we estimate pure forecasting regressions, replacing $\text{Crisis}_{i,t \text{ to } t+2}$ with $\text{Crisis}_{i,t+1}$

Table 2: *Crisis prediction models*

	OLS			IV		
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta_3\text{Rate}_t$	0.02^{***}	0.02^{***}	0.01^{**}	0.03^{**}	0.03^{**}	0.00
	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)
$\text{Cut Rate}_{t-8,t-3}$		0.06^{***}	0.07^{***}		0.06^{**}	0.06^{***}
		(0.02)	(0.02)		(0.02)	(0.02)
$\Delta_3\text{Rate}_t \times \text{Cut Rate}_{t-8,t-3}$			0.03^{***}			0.07^{***}
			(0.01)			(0.03)
Country fixed effects	✓	✓	✓	✓	✓	✓
Kleibergen-Paap Weak ID				93.83	95.95	42.94
Observations	1625	1625	1625	1625	1625	1625

Notes: This table shows linear probability models for a systemic banking crisis occurring between years t and $t+2$. All specifications control for 8 lags of GDP growth and inflation. IV specifications instrument policy rate changes with the residualized Jordà et al. (2020) trilemma variable. Δ_3Rate is the 3-year change in the nominal monetary policy rate. Cut is a dummy which equals 1 if nominal rates were cut between $t-8$ and $t-3$. IV interaction specifications include residualized JST trilemma variable and its interaction with the cut dummy as instruments. In this case the Kleibergen-Paap Weak ID is the joint test for both instruments. Country-clustered standard errors in parentheses. *, **, and *** indicate significance at the 0.1, 0.05, and 0.01 levels, respectively.

country fixed effects and lags of GDP growth and inflation as controls. Similar to the results in Schularick et al. (2021), we find that an increase in policy rates is associated with elevated crisis risk over the following years. A 1 percentage point cumulative increase in interest rates between years $t-3$ and t is associated with a 2 percentage points higher probability of a systemic banking crisis between years t and $t+2$. Column 2 shows that a cut in rates between $t-8$ and $t-3$ predicts elevated crisis risk: the probability of a crisis between years t and $t+2$ increases by 6 percentage points if rates were cut between $t-8$ and $t-3$. In column 3, we additionally include the interaction between the two variables, corresponding to equation (2). The interaction coefficient is highly significant, suggesting that increases in policy rates predict crisis risk particularly when there has been a cut in rates previously. A 1 percentage point cumulative increase in rates during years $t-3$ to t is associated with 4 percentage points higher crisis probability during years t to $t+2$ if rates were previously cut (sum of Δ_3Rate and $\Delta_3\text{Rate} \times \text{Cut}$ coefficients in column 6), and only 1 percentage point higher crisis probability if they were not (the Δ_3Rate coefficient).

In Table 2 columns 4 to 6 we instrument three year changes in policy rates (Δ_3Rate) with the residualized trilemma shocks from Jordà et al. (2020), coded according to equation (3). Instrumented three-year changes in policy rates in column 4 are associated with higher crisis likelihood. The first-stage relationship between the trilemma instrument and policy rates is strong, as indicated by the Kleibergen-Paap Weak ID statistic reported at the bottom of the table. In column 6 we additionally include the interaction of the three-year changes in policy rates and the dummy variable for a previous cut from $t-8$ to $t-3$, instrumenting the rate change with the change in the

Table 3: *Crisis prediction: real interest rates and inflation*

	Real rates			Inflation		
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta_3\text{Var}_t$	0.003*	0.002	0.002	-0.001	-0.001	-0.001
	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)
$1(\Delta\text{Var}_{t-8,t-3} < 0)$		0.013	0.013		-0.013	-0.013
		(0.020)	(0.020)		(0.020)	(0.020)
$\Delta_3\text{Var}_t \times 1(\Delta\text{Var}_{t-8,t-3} < 0)$			0.001			0.002
			(0.003)			(0.002)
Country fixed effects	✓	✓	✓	✓	✓	✓
Observations	1942	1942	1942	2104	2104	2104

Notes: This table shows linear probability models for a systemic banking crisis occurring between years t and $t+2$. Var refers to the real interest rate in columns 1–3, and inflation in columns 4–6. Inflation is the change in the Consumer Price Index. Real interest rate is the difference between the nominal rate and inflation in the same year. OLS regressions with country fixed effects. Country-clustered standard errors in parentheses. *, **, and *** indicate significance at the 0.1, 0.05, and 0.01 levels, respectively.

trilemma IV, and the interaction with the change in trilemma IV interacted with the cut dummy. The Kleibergen-Paap weak ID test reported below is the joint test for both instrumental variables. The results confirm that the association between rate hikes and crisis risk is primarily driven by those rate hikes which happen after a period of rate cuts as indicated by the highly significant interaction term.

Looking at Table 2 column 6, a rate increase between years $t-3$ and t is not associated with higher crisis risk if rates were previously raised (Cut dummy equals zero): the coefficient on Δ_3Rate is zero and statistically insignificant. A rate increase after a series of cuts, on the contrary, is associated with a 7 percentage point higher crisis probability for every 1 percentage point cumulative 3-year increase in monetary policy rates. Summing together the Cut and $\Delta_3\text{Rate} \times \text{Cut}$ coefficients in column 6, a sequence of reducing rates and then increasing them by 1 percentage point over three years is associated with a 13 percentage points increase in crisis risk, more than doubling the crisis probability compared to the sample mean of 10% (given an annual crisis probability of around 3.5%, we have a probability of 10% for at least one crisis observation in a three-year window). Together, this elevated crisis probability of around 23% (10% sample mean plus 13% increase driven by the cut and then increase in rates) aligns well with the 20% crisis frequency after U-shaped monetary policy reported in Table 1.

In Appendix Table A.6, we show that the relationship in Table 2 holds for 1-year ahead crisis prediction, in post-1945 data, using Driscoll-Kraay standard errors and including decade fixed effects and a large set of control variables. Appendix Table A.7 additionally shows that results hold if we use probit rather than linear probability models of crisis risk. Finally, Table A.8 shows results for a different functional form of the path of policy rates. We first measure the length (in years) of a spell of low interest rates (low in year t refers to below moving average over the preceding 10-year window). The table then shows that the interaction between rate changes and the length of the preceding low interest environment is a significant predictor of crises.

Inflation and real rates We next compare the results on crisis prediction with nominal rates to real rates and inflation. Unfortunately, in the case of real rates and inflation there is no instrumental variable to rely on. Hence, we report the OLS results for both of them in Table 3. Different to nominal interest rates, higher real rates are not robustly associated with higher crisis risk, and there are no clear state-dependencies concerning the real interest rate path and risk of future crises in the data. A similar result holds for inflation, shown in columns 4 to 6, with increases in inflation not associated with a higher risk of crises, and no path-dependencies between the sequence of high and low inflation episodes on the one hand, and crisis risk on the other.

4. Understanding the mechanisms in macro data

Why is the U shape in policy rates associated with higher crisis risk? One reason may be that decreasing rates in years $t-8$ to $t-3$ before the crisis encourages excessive lending and asset price growth, generating financial vulnerabilities. The subsequent rate increases could then force these asset price and credit booms to unwind, crystallizing the corresponding vulnerabilities.

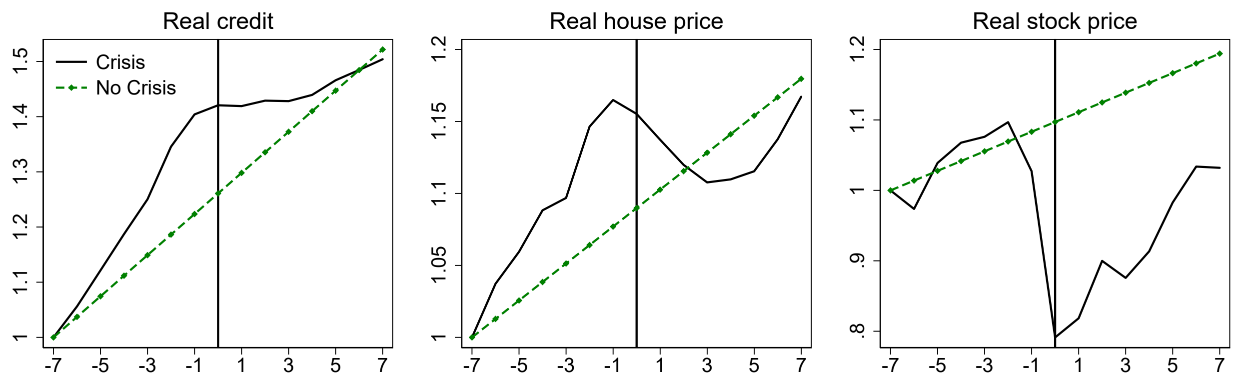
Figure 4a confirms that crises are accompanied by a boom-bust cycle in credit and asset prices, in line with findings in previous studies (Schularick and Taylor, 2012; Krishnamurthy and Muir, 2017; Mian, Sufi, and Verner, 2017). It shows the cumulative growth rates in real credit, house prices, and stock prices from 7 years before to 7 years after the crisis start date. These windows are constructed by averaging the growth rate of the corresponding variable for a given pre- or post-crisis year, and cumulating these growth rates starting with 1 at $t-7$. Credit and house prices grow at above-average levels before the crisis, and grow at below-average levels in the post-crisis years. For stock prices, there is less of a pre-crisis boom, but the crisis is accompanied by sharp stock price decline which take some years to normalize.

Figure 4b shows that this boom-bust pattern in credit and asset prices is accompanied by a cycle in economic activity, with both real GDP and consumption growing at above average rates during the pre-crisis boom, and at below-average rates during the post-crisis bust. This boom-bust cycle is particularly pronounced for a measure of real activity that is likely to be particularly sensitive to credit conditions and house prices: housing investment. The cumulative growth rates in this variable are some 25 percentage points above trend before the crisis, followed by a sharp reversal in the form of housing investment declines during the post-crisis years. Appendix Figure A.4 shows that similar boom-bust patterns in asset prices and real activity are observed for other measures of systemic banking crises.

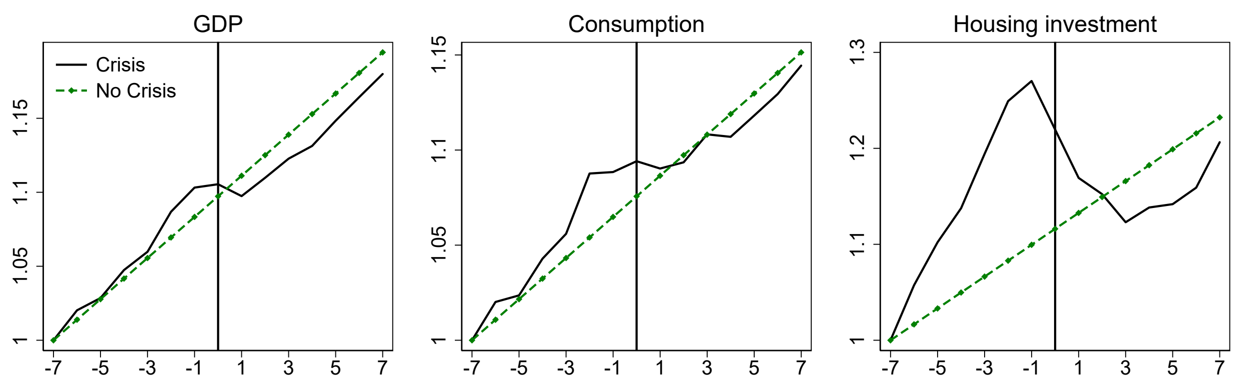
Does the path of monetary policy make a contribution to these booms and busts in credit, asset prices, and real activity? To assess this, we test whether cuts in monetary policy rates contribute

Figure 4: *Credit, asset prices, and real activity around crises*

(a) *Credit and asset prices*



(b) *Real activity*



Notes: Unweighted averages of 72 systemic banking crises defined as in Jordà et al. (2016). To construct the figures, we calculate average log growth rates of each variable in a given pre- or post-crisis year, and cumulate them starting at 1, from 7 years before to 7 years after the crisis start date.

to the growth of credit and asset prices during the boom, and if subsequent policy rate increases manifest themselves in a particularly strong adjustment in these financial and real variables.

4.1. Monetary policy and the pre-crisis boom

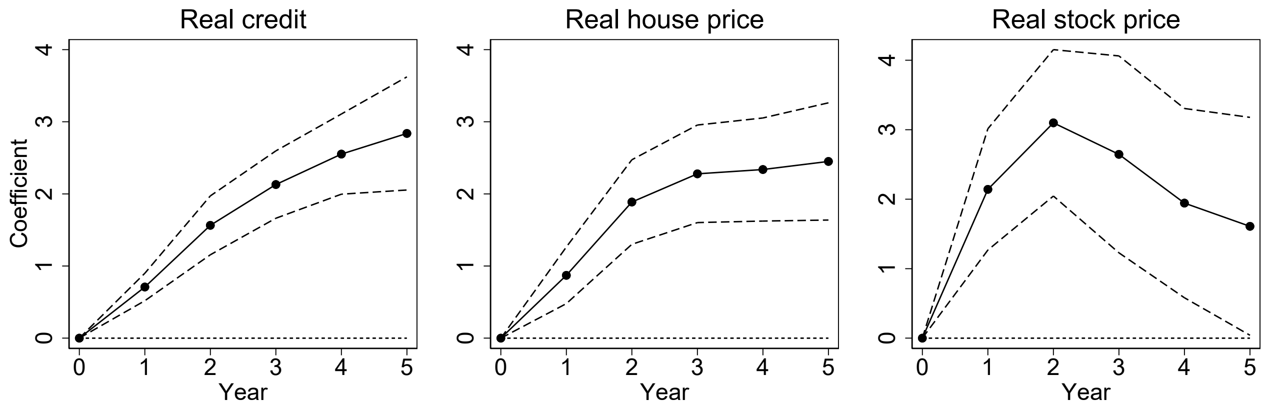
Rate cuts and booms in credit and asset prices Can loose monetary policy fuel booms in credit and asset prices? To study this question, we run the following local projection (Jordà, 2005):

$$\Delta_h y_{i,t\to t+h} = \alpha_{i,h} + \alpha_{d,h} + \beta_h \Delta\text{Rate}_{i,t} + \sum_{L=0}^{L=4} \gamma_l X_{i,t-L} + \epsilon_{i,t+h}, \quad h \in \{1, ..., 5\}. \tag{4}$$

Above, y is the log growth in real credit, real house price, or real stock price, between years t and $t+h$, for horizons h 1 to 5 years ahead. ΔRate is the change in the policy rate, raw or instrumented using the Jordà et al. (2020) trilemma instrument. X are control variables (four lags and contemporaneous) consisting of real credit, house, and stock price growth, short and long-term

Figure 5: *Response of credit and asset prices to policy rate cuts*

(a) *OLS:*



(b) *IV:*

Notes: Local projection estimates of future real credit, house price, and stock price growth, for horizons $t+1$ to $t+h$ on the change in policy rates at t, raw (panel a) and instrumented using the Jordà et al. (2020) trilemma IV (panel b). All regressions include country and decade fixed effects, and control for contemporaneous levels and four lags of real credit, house price, stock price, GDP, consumption, and housing investment growth, as well as four lags of changes in short- and long-term rates. Dashed lines show 90% confidence intervals.

rate changes, as well as GDP, consumption, and housing investment growth. α_i and α_d are country and decade fixed effects. Standard errors ϵ are clustered by country. Because we are interested in responses to policy during the boom, we exclude the 3 years immediately after the start of a crisis from our sample.

Figure 5 shows the responses of credit, house prices, and stock prices to policy rate innovations, measured as the coefficients β_h in equation (4). The coefficients are standardized to a 1ppt reduction in rates. Figure 5a shows these responses for raw cuts in policy rates, and Figure 5b shows the results for rate changes instrumented using the Jordà et al. (2020) trilemma IV. In line with previous studies, we find that loose monetary policy is followed by faster than usual growth in asset prices and credit. These effects are statistically significant and economically large: after 5 years, a 1 percentage point policy rate cut leads to between 6 and 15 percentage points higher cumulative real credit and asset price growth in the IV specifications.

Appendix Figure A.5 shows that the patterns documented in Figure 5 hold if we restrict the sample to the post World War 2 period, and if we restrict our attention to large cuts in policy rates (i.e., instrumented interest rate changes that are in the lowest quartile of the distribution), instead of treating policy rate cuts and increases symmetrically. Finally, Appendix Figure A.6 shows that rate cuts are followed by large increases in both household mortgage and business loans, and falls risk premia on real estate and business equity measured, correspondingly, as the rent-price and dividend-price ratios. These results are consistent with the idea that cuts in monetary policy rates lead to increases in the supply of credit to both households and businesses.

Decreases in policy rates are associated on average with surging asset prices and high credit growth over the next few years. In this section, we want to ask whether we can link decreases in policy rates to a joint measure of pre-crisis vulnerabilities which captures both the boom in credit, and the surge in asset prices. To do this, we rely on a simple proxy recently established by Greenwood et al. (2022), the financial "red zone" (R-zone). To study periods when asset prices and credit are jointly elevated, they define a country to be in the R-zone if three-year credit growth and asset price growth are both elevated relative to their respective sample averages. More specifically, we follow them and define the R-zone indicator variable as follows

$$
\begin{aligned}
\text{R-zone}_{i,j,t} &= \text{High-Credit-Growth}_{i,j,t} * \text{High-Price-Growth}_{i,j,t}, \quad \text{where} \\
\text{High-Credit-Growth}_{i,j,t} &= 1\left\{\Delta_3(\text{Credit/GDP})_{i,j,t} > 80^{\text{th}} \text{ percentile}\right\}, \quad \text{and} \\
\text{High-Price-Growth}_{i,j,t} &= 1\left\{\Delta_3 \ln(\text{Asset Price})_{i,j,t} > 66.7^{\text{th}} \text{ percentile}\right\}.
\end{aligned}
$$

Above, i, j, and t are country, economic sector, and year indices, credit refers to total credit to the business or household sector, and asset price refers to equity (for the business sector) and house (for the household sector) prices deflated by CPI. If no decomposition into household and business credit is available, we define the high-credit-growth dummy based on total private credit. The R-zone indicator refers to joint high growth in credit and asset prices, and can be defined separately for the specific sector (business or household), or jointly based on either of these two sectors being in the R-zone.

Rate cuts and the R-zone Based on this chronology, we then ask whether ending up in the R-zone is predicted by the past path of policy rates. To this end, we rely on the simple rate cut indicator variable $\text{Cut}_{t-5,t}$ introduced in Section 3.2 and equation (2), and ask whether cuts in policy rates over the past 5 years are associated with the probability of being in the R-zone over the following three years. The dependent variable is hence a three-year moving average of the R-zone indicator over the years $t+1$ to $t+3$. Table 4 shows that decreasing policy rates are associated with a higher likelihood of ending up in the business or household R-zone. Appendix Table A.9 confirms that R-zone events, as in

Table 4: *Rate cuts and the financial "red zone"*

	R-Zone Business$_{t+1 \text{ to } t+3}$ (1)	R-Zone Households$_{t+1 \text{ to } t+3}$ (2)	R-Zone Either$_{t+1 \text{ to } t+3}$ (3)
Cut Rate$_{t-5,t}$	0.12***	0.28***	0.30***
	(0.04)	(0.06)	(0.06)
Country fixed effects	✓	✓	✓
Controls	✓	✓	✓
Observations	1650	1475	1712

Notes: Red zone ("R-zone") is defined as joint high growth in credit and asset prices, using the same thresholds as in Greenwood et al. (2022). We use business credit and equity prices for the business R-zone, and household credit and house prices for the household R-zone. We use total credit growth as proxy for business or household credit growth when the business/household split is not available. All regressions control for five lags of real GDP growth and inflation. Country-clustered standard errors in parentheses. *, **, and *** indicate significance at the 0.1, 0.05, and 0.01 levels, respectively.

Greenwood et al. (2022), are associated with a higher likelihood of crisis.

4.2. Policy rate reversal and macro-financial adjustment

The results in section 4.1 suggest that rate cuts over a long time period (in our regression specification, 5 years) lead to booms in credit and asset prices, which in turn bring the economy into a vulnerable R-zone, in turn helping explain why crises become more likely after rates are cut. We next examine why, after a series of such rate cuts takes place, raising rates can crystallize the associated financial vulnerabilities and trigger a crisis. To do this, we test whether lending and asset prices become more sensitive to changes in interest rates after a series of rate cuts, using the same method as we used to test for rate cut state-dependencies in crisis prediction in Section 3.2 and equation (2). More specifically, we run the following regression:

$$\Delta_3 y_{i,t-1\to t+h} = \alpha_i + \beta_1 \Delta_3 \text{Rate}_{t-3,t} + \beta_2 \text{Cut}_{t-8,t-3} + \beta_3 \Delta_3 \text{Rate}_{t-3,t} \times Cut_{t-8,t-3} + \sum_{L=1}^{L=4} \gamma_l X_{i,t-L} + u_{i,t-1\to t+h}, \quad (5)$$

Above, y is the cumulative log growth in credit or asset prices between year $t-1$ and $t+h$, where we will focus on results for $h=2$ and $h=5$. Cut is a dummy variable indicating whether rates were cut between years $t-8$ and $t-3$, and Rate is the change in the policy rate between years $t-3$ and t, in percentage points, instrumented with the Jordà et al. (2020) trilemma instrument. X is a vector of control variables that are the same as in the local projection equation (4) in Section 4.1 (four lags of real credit, house price, stock price, GDP, consumption, and housing investment growth), plus four lags of the crisis indicator to

Table 5: *Policy rate reversal, credit, and asset prices*

	Loans		House Prices		Stock Prices	
	3-yr (1)	6-yr (2)	3-yr (3)	6-yr (4)	3-yr (5)	6-yr (6)
$\Delta_3\text{Rate}_t$	-0.97*	-2.75***	-2.51***	-2.00**	-2.53*	-7.33***
	(0.56)	(0.71)	(0.82)	(0.85)	(1.52)	(2.05)
Cut $\text{Rate}_{t-8,t-3}$	-0.02*	-0.03	0.02	0.02	-0.07	0.04
	(0.01)	(0.02)	(0.01)	(0.02)	(0.05)	(0.04)
$\Delta_3\text{Rate}_t \times$ Cut $\text{Rate}_{t-8,t-3}$	-1.94***	-3.00**	-1.23	-2.33**	-7.24***	0.52
	(0.61)	(1.19)	(0.93)	(1.12)	(1.99)	(1.89)
Kleibergen-Paap Weak ID	18.49	18.49	19.75	19.75	18.65	18.61
Observations	1054	1054	1061	1061	1056	1060

Notes: y variables in columns refer to cumulative log growth in the real variable (e.g., loans/CPI) between $t-1$ and $t+2$ (3-yr specifications), and $t-1$ and $t+5$ (6-yr specifications). x variable is the policy rate change instrumented by the Jordà et al. (2020) trilemma IV. All regressions include country and decade fixed effects, and control for four lags of real loan, house price, stock price, GDP, consumption, and housing investment growth, and four lags of the crisis dummy. Country-clustered standard errors in parentheses. *, **, and *** indicate significance at the 0.1, 0.05, and 0.01 levels, respectively.

control for whether rate cuts starting in year $t-3$ had already triggered a crisis. A negative coefficient $\beta_1 < 0$ will tell us that credit and asset prices respond negatively to increases in policy rates. A coefficient $\beta_3 < 0$ will tell us that these responses become stronger if these rate increases were preceded by a series of rate cuts.

The results of these regressions are shown in Table 5. Columns 1, 3, and 5 show the coefficients for the 3-year ahead horizon (growth from $t-1$ to $t+2$), the same as in the crisis prediction regressions in Table 2), and columns 2, 4, and 6 show the coefficients for a longer 6-year horizon (growth from $t-1$ to $t+5$). The negative and significant coefficient on the Δ_3Rate variable (top row) shows that rate increases lead to falls in credit and asset prices. The Cut dummy alone is not robustly associated with weaker credit and asset price growth (middle row). But when the rate cuts are followed by an increase, the subsequent declines in credit and asset prices are much larger than if these rate increases were not preceded by the cuts, as indicated by the large negative signs on the interaction coefficient in the bottom row of Table 5.

The magnitudes of these state-dependencies are statistically significant and economically large. A 1 percentage point policy rate increase alone leads to 1–2.5 percentage points declines in credit and asset prices over a three-year window. But if this increase was preceded by a series of cuts, the effects on credit and asset prices become 2–4 times stronger, with between 3 and 10 percentage point declines 3 years ahead (measured as the sum of the coefficients on Δ_3Rate and the Δ_3Rate x Cut interaction). These state-dependencies are always present for real credit growth, but for house and stock prices, their importance

Table 6: *Financial red zone, rate hikes, and crises*

	Dependent variable: $\text{Crisis}_{t \text{ to } t+2}$					
	R-Zone Business		R-Zone Households		R-Zone Either	
	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)
$\text{R-Zone}_{t-3 \text{ to } t-1}$	0.07^{***}	0.06^{***}	0.10^{***}	0.09^{***}	0.06^{***}	0.06^{***}
	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
$\text{R-Zone}_{t-3 \text{ to } t-1} \times \Delta_3\text{Rate}_t$	2.47^{**}	4.33^{**}	3.00^{***}	3.23^{*}	2.56^{***}	3.17^{**}
	(1.17)	(1.74)	(0.87)	(1.82)	(0.71)	(1.51)
$\Delta_3\text{Rate}_t$	1.24^{***}	1.43	0.64^{*}	1.12	0.60^{**}	0.58
	(0.29)	(1.21)	(0.34)	(1.12)	(0.26)	(1.20)
Country fixed effects	✓	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓	✓
Kleibergen-Paap Weak ID		38.62		28.00		32.77
Observations	1659	1369	1478	1235	1722	1427

Notes: Financial red zone ("R-zone") is defined as joint high growth in credit and asset prices, using the same thresholds as in Greenwood et al. (2022). We use business credit and equity prices for the business R-zone, and household credit and house prices for the household R-zone. We use total credit growth as proxy for business or household credit growth when the business/household split is not available. All regressions control for five lags of real GDP growth and inflation. Country-clustered standard errors in parentheses. *, **, and *** indicate significance at the 0.1, 0.05, and 0.01 levels, respectively.

depends on the horizon. Comparing Table 5 columns 3 and 4 to columns 5 and 6, house prices react more strongly to the rate cut preceded by a raise at longer horizons of 6 years, while stock prices react more strongly at the shorter horizon of 3 years. This is in line with the crisis window averages shown in Figure 4, which, after the crisis, show large immediate declines in stock prices followed by a recovery, and a protracted slow decline in house prices. Appendix Table A.10 shows that these results hold if we run simple OLS as opposed to an IV regression.

Raising rates in the R-zone Greenwood et al. (2022) suggested that in order to deflate the booms in credit and asset prices, the policymakers could lean against the wind by raising interest rates. Based on the arguments presented in our paper so far, however, once in the R-zone, such a policy action could instead trigger the crisis by crystallizing the financial vulnerabilities that built up during the period of low interest rates. We next test directly whether raising rates in the R-zone, and bursting the credit and asset price booms as indicated by the results in Table 5, can trigger rather than prevent a financial crisis. To do this, we predict financial crises using changes in policy rates, and compare the magnitudes of these effects conditional on being or not being in the R-zone.

As in our baseline specification in Section 3.2, we estimate the likelihood of experiencing a crisis over a three-year window as a function of three-year changes in interest rates, now including in addition a measure of being in the R-zone over the previous three years, and its interaction with changes in interest rates. The results of these regressions are shown in Table 6. We find that the probability of banking crises is significantly elevated when countries are in the R-zone and policy rates are increased. These results hold in OLS and IV specifications, where we instrument three-year changes in interest rates and their interaction with the R-zone indicators with the cumulative trilemma instrument and its interaction with R-zone dummies and report the joint Kleibergen-Paap test.

The corresponding economic magnitudes are large. Focusing on the IV specifications in Table 6 columns 2, 4, and 6, raising rates outside of the R zone is not significantly associated with a higher risk of crises. This mirrors our results in Table 2 (that raising rates after a series of rate increases does not increase crisis risk), and suggests that if the policymakers act early enough, before the economy is in the R-zone, it is possible to deflate the asset price booms without triggering a crisis (since the coefficients on Δ_3Rate alone in Table 5 are negative and significant). If policymakers raise rates when the economy is already in the R-zone, however, a crisis is much more likely. Summing the coefficients on R-zone and the interaction between Δ_3Rate and the R-zone indicator, raising rates by 1 percentage point when the economy is in the R-zone increases the probability of a financial crisis by 9–12 percentage points over the next 3 years. This amounts to roughly doubling the unconditional probability of a crisis (10% in the full sample), with the sum of the unconditional probability and the additional effect of raising rates in the R-zone, again, aligning well with the 20% crisis frequency after U-shaped monetary policy in Table 1.

Policy rate reversal and macroeconomic outcomes Do these credit and asset price cycles—first inflated, and then deflated by monetary policy—have real effects? To test for this, we run the same regression as in equation (5), but put measures of real activity on the left-hand side. More precisely, we regress 3-year real growth rate in three measures of real activity—GDP, consumption, and housing investment—on changes in the policy rate, and changes in rates interacted with the Cut dummy, as well as the standalone Cut dummy and controls. This allows us to directly test whether U-shaped monetary policy is associated with poor macroeconomic outcomes, as well as measures of financial stress.

The results are reported in Table 7. In line with previous studies, we find that policy rate increases lead to declines in GDP and consumption (columns 1–4), and we find a similar result, albeit much stronger in magnitude, for housing investment (columns 5–6). In addition to this, we also find a clear state dependency similar to that found for financial variables: rate increases which were preceded by cuts lead to much larger declines in these

Table 7: *Policy rate reversal and real macroeconomic outcomes*

	GDP		Consumption		Housing investment	
	3-yr (1)	6-yr (2)	3-yr (3)	6-yr (4)	3-yr (5)	6-yr (6)
$\Delta_3\text{Rate}_t$	-0.69***	-1.03***	-0.47*	-0.88**	-3.93***	-3.76**
	(0.26)	(0.35)	(0.24)	(0.43)	(1.36)	(1.58)
Cut $\text{Rate}_{t-8,t-3}$	-0.00	0.01	-0.00	0.01	-0.01	-0.00
	(0.01)	(0.01)	(0.01)	(0.01)	(0.03)	(0.01)
$\Delta_3\text{Rate}_t \times$ Cut $\text{Rate}_{t-8,t-3}$	-0.88***	-1.13**	-0.72**	-1.26**	-3.43*	-3.69**
	(0.29)	(0.49)	(0.34)	(0.55)	(1.90)	(1.84)
Kleibergen-Paap Weak ID	20.00	20.00	20.00	20.00	19.43	19.43
Observations	1066	1066	1066	1066	1045	1045

Notes: y variables in columns refer to growth in the real variable (e.g., real GDP) between $t-1$ and $t+2$ (3-yr specifications), and $t-1$ and $t+5$ (6-yr specifications). x variable is the policy rate change instrumented by the Jordà et al. (2020) trilemma IV. All regressions include country and decade fixed effects, and control for four lags of real loan, house price, stock price, GDP, consumption, and housing investment growth, and four lags of the crisis dummy. Country-clustered standard errors in parentheses. *, **, and *** indicate significance at the 0.1, 0.05, and 0.01 levels, respectively.

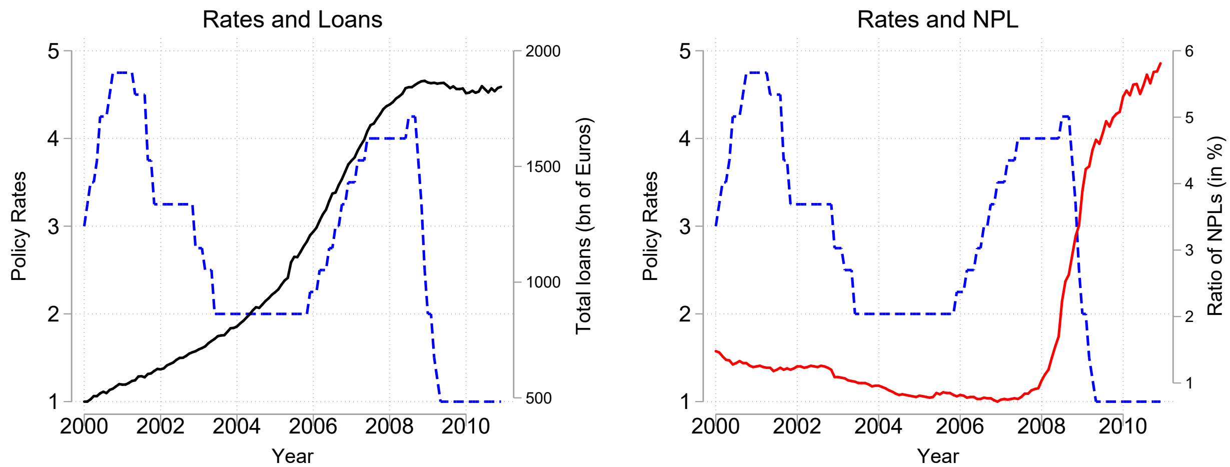
three measures of real activity. Comparing the coefficient on Δ_3Rate with the sum of the Δ_3Rate coefficient and the interaction Δ_3Rate x Cut Rate coefficient, the effect of monetary policy and GDP, consumption, and housing investment is roughly two times stronger when the policy rate increase was preceded by a series of cuts, both at the 3-year and 6-year horizons. For example, Table 7 column 1 shows that when rates were previously increased, raising policy rates by 1 percentage point leads to a 0.7 percentage points lower cumulative real GDP growth over the next 3 years. When rates were previously cut, the reduction in real GDP growth is, instead, 1.7 percentage points (the sum of the coefficients on Δ_3Rate, 0.7, and the interaction $\Delta_3\text{Rate}\times\text{Cut}$, 0.9). Appendix Table A.11 shows that these results hold for simple OLS, as well as the IV in Table 7.

Taken together, our findings suggest that monetary policy is an important driver not only of the crisis boom-bust pattern in credit and asset prices shown in Figure 4a, but also the boom-bust in real activity that takes place around crises, shown in Figure 4b.

5. Loan-level evidence from the Spanish credit register

To study the potential mechanisms in more detail, we now turn to an in-depth analysis of monetary cycles and credit risk in Spain. Spain has three qualities that make it an ideal candidate for our empirical strategy. First, it has a financial system dominated by banks, which means that studying the provision of bank credit in detail gives us a good overview of the overall financial developments in the country. Second, it has a fairly exogenous

Figure 6: *Monetary policy rates, credit growth, and loan defaults in Spain around the 2007–08 crisis*



Notes: The left panel shows the interest rate for main refinancing operations (in percent, dashed blue line), and total loans to the private domestic sector (€billions, solid black line) in Spain for years 2000 to 2010. The right panel shows the interest rate (dashed blue line) alongside the ratio of non-performing loans to total loans (solid red line). Data are monthly (interest rates are annualized).

monetary policy. Spain was part of the European Exchange Rate Mechanism (ERM) in the 1990s, and joined the euro in 1999. This means that monetary policy during the time period of our sample, 1995–2020, closely mirrors our identification strategy in the aggregate, with Spanish monetary policy rates decided in Frankfurt with a focus on the core countries of the euro. Third, Spain has granular data at loan level that covers a long period of time. These features will allow us to understand how credit risk is associated with monetary policy rates at the level of an individual loan. More specifically, we will use the micro data to, first, validate the correlations that we established in the macroeconomic data, and, second, explore the heterogeneous responses at bank and firm level to better understand their underlying drivers.

To provide some background, Figure 6 shows that the aggregate developments in Spain during the pre-crisis boom of the early 2000s mirrored the average patterns for historical crises shown in Figures 2 and 4. The left panel shows that monetary policy rates (dashed blue line) first decreased, hitting their lowest point of 2% in years 2004–05. The decisions by the European Central Bank to cut interest rates mainly corresponded to the need to boost economic growth in core European countries such as Germany, and did not reflect the economic situation in Spain which, on the contrary, experienced high GDP growth and inflation during this time period. Especially for this period, monetary policy rates in Spain can be viewed as exogenous, while policy in the 2010s was more aligned to economic developments in the euro-area periphery countries.

During the time of low interest rates, credit expansion accelerated and credit grew at a higher rate than in the pre-2004 environment. After a series of policy rate hikes up to 4% credit growth slowed down during 2007–08. At the same time, the right panel shows that non-performing loans as a share of total loans declined in the low interest environment, and started to increase rapidly as early as the beginning of 2008, when policy interest rates peaked. Over the twelve months from August 2007 to August 2008, when interest rates had already increased, but before Lehman went bankrupt, the ratio of non-performing loans to total loans tripled from 0.8% to 2.5%. After the Lehman default, this ratio increased further, reaching 6% in 2010.

Sources of administrative data We will now study the drivers of these developments at the level of individual loan. Banco de España, in its role as bank supervisor of the Spanish financial system, is the owner of the Spanish Credit Register (CIR), a confidential credit bureau which collects detailed monthly information on new and outstanding loans to firms (with amounts exceeding 6,000 euros), granted by all credit institutions operating in Spain since 1984, and defaults of these loans. This dataset not only contains information about the loan (e.g., size, maturity, level of collateralization), but also on the identity of the bank that grants the loan and on the identity of the firm that receives it, which gives us the opportunity to expand the original dataset with bank balance sheet information (also compiled by Banco de España at monthly frequency), and firm economic and financial information from the Spain's Mercantile Register (collected annually since 1995).

For the purposes of the paper we, therefore, work with quarterly data from 1995 to 2020 containing newly granted loans to non-financial companies by commercial banks, savings banks, and credit cooperatives. These financial institutions account for more than 95 percent of bank debt in the Spanish financial system. We analyze these data at the firm-bank level. To this end, we aggregate all new loans granted by a given bank to a given firm in the same quarter. Moreover, to keep the data manageable, we take a 10% random sample of the entire population of Spanish firms. Based on the sample of all newly granted loans to these firms, we study the determinants of loan defaults at one year ahead horizon. We define default, or delinquency, as the loan payments being more than 90 days overdue, following the main definition used by policymakers. Appendix Table A.12 provides the summary statistics (mean, standard deviation, quartiles) of the variables used in the analysis.

Empirical specifications We analyze the link between ex-post loan defaults and the shape of monetary policy, controlling for a large set of observable and unobservable firm, bank and macro variables. We start by analyzing average effects to confirm that the findings from Sections 2–4 hold in the micro data. We then turn to study firm and bank level heterogeneities to gain additional insights into the forces driving these average effects in macro and micro data. Specifically, we estimate by OLS the following linear probability model:

$$\text{Loan Default}_{i,j,t,t+1} = \beta_1 \Delta\text{Rate}_{t,t+1} + \beta_2 \text{Cut}_{t-5,t} + \beta_3 \Delta\text{Rate}_{t,t+1} \times \text{Cut}_{t-5,t} + \gamma_1 F_{i,t-1} + \gamma_2 B_{j,t-1} + \gamma_3 M_t + u_{i,j,t,t+1} \quad (6)$$

where i refers to the firm, j to the bank that grants the loan and t to time. Loan Default is a dummy variable that takes the value of one if the loan granted by bank j at time t to firm i becomes delinquent in the subsequent year, and 0 otherwise.[5] ΔRate is the change in the overnight interest rate between years t and $t+1$, in percentage points. Cut is a dummy variable indicating whether the change in overnight interest rates between years $t-5$ and t is below its average value.[6] The specification closely mirrors our approach in the aggregate data, focusing on loans that were granted during a period of decreasing monetary policy rates, and studying their default risk when rates start increasing.

We saturate these specifications with a large set of control variables and fixed effects. F is a vector of control variables and fixed effects at firm level that includes, depending on the specification, industry (NACE at 3 digits) and location (zip code level) fixed effects, firm fixed effects, firm$\times$bank fixed effects, and a set of observable characteristics (in the previous year) such as log of assets, log of age, own funds over total assets, liquid assets over total assets, ROA, bad credit history (a dummy that takes value of 1 if the firm has defaulted before time t, and 0 otherwise) and the firm's average cost of credit. B is a vector of control variables and fixed effects at bank level that includes bank fixed effects and/or observable characteristics (in the previous quarter) such as log of total assets, capital and liquidity ratios, ROA and NPL ratios. Finally, M_t is a vector of macro controls and fixed effects that includes GDP and CPI in the same structure as interest rates, and—when heterogeneous effects are tested—it also includes time fixed effects. u is the error term. Standard errors are multi-clustered at the firm, bank and time level to allow for serial correlation across firms, banks and time periods.

The β_1 coefficient reflects the relationship between loan defaults and changes in policy rates after loan origination. A positive β_2 coefficient means that the probability of loan default is higher when the loan is granted in periods of loose monetary policy. The coefficient on the interaction term, β_3, is testing for the role of U-shaped monetary policy:

[5]We want to analyze defaults for loans granted in periods of low monetary rates and what happens to defaults when monetary policy rates increase. Because half of the new loans have a maturity shorter than 1 year (see Table A.12 in the Appendix), we analyze one-year probability of loan default.

[6]Since monetary policy rates were on average decreasing in our sample (see Appendix Table A.12), the Cut dummy variable effectively corresponds to large cuts in interest rates.

Table 8: *The path of monetary policy and loan-level defaults in Spain*

	Dependent variable: Loan default at t+1								
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$\Delta \text{Rate}_{t,t+1}$	0.002^{***}	0.002^{**}	0.002^{**}	0.002^{**}	0.002^{***}	0.003^{***}	0.002^{*}	0.001^{*}	0.001^{*}
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Cut $\text{Rate}_{t-5,t}$	0.005^{***}	0.007^{***}	0.004^{***}	0.004^{***}	0.003^{**}	0.003^{**}	0.002^{*}	0.004^{***}	0.004^{***}
	(0.001)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
$\Delta \text{Rate}_{t,t+1} \times$ Cut $\text{Rate}_{t-5,t}$							0.002^{**}		0.002^{***}
							(0.001)		(0.001)
Industry×Location FE	No	Yes	Yes	Yes	-	-	-	-	-
Bank Controls	No	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Bank FE	No	No	No	Yes	Yes	-	-	-	-
Firm FE	No	No	No	No	Yes	-	-	-	-
Firm×Bank FE	No	No	No	No	No	Yes	Yes	Yes	Yes
Firm Controls	No	No	No	No	No	No	No	Yes	Yes
Observations	1.75m	1.75m	1.75m	1.75m	1.75m	1.75m	1.75m	1.27m	1.27m
R^2	0.016	0.098	0.098	0.099	0.183	0.318	0.318	0.351	0.351

Notes: This table reports the OLS regression results of the probability that a loan granted at time t becomes delinquent in the next year (t+1) for different paths of monetary policy rates. Cut is a dummy variable indicating whether the change in (overnight) rates between years t-5 and t is below its average value, and ΔRate is the percentage point change in the policy rate between years t and t+1. Coefficients are listed in the first row, and standard errors corrected for clustering at the firm, bank and time level are reported in the row below. "Yes" indicates that the set of characteristics or fixed effects is included, "No" that is not included and "-" that is comprised by the included set of fixed effects. For observations, m corresponds to millions. *, **, and *** indicate significance at the 0.1, 0.05, and 0.01 levels, respectively.

a positive β_3 coefficient means that the relationship between loan defaults and changes in monetary policy rates becomes stronger if the policy rate changes were preceded by a period of declining interest rates.

Results Table 8 reports results of the regressions specified in equation (6). We start with a model free of controls to progressively saturate the specification with observable and unobservable bank and firm characteristics. In column 1 there are no controls, and both the coefficients on Cut and ΔRate are positive and statistically significant. This tells us that 1-year ahead loan defaults increase when interest rates decreased in the 5 years before the loan was granted, and when there is an increase in policy rates after loan origination. In column 2 we include industry×location fixed effects, in column 3 we include bank controls, in column 4 bank fixed effects, in column 5 firm fixed effects and in column 6 firm×bank fixed effects. In all these specifications we obtain very similar results, although the R^2 increases from 0.016 to 0.318, suggesting that the estimated effects do not suffer from biases due to further unobservable omitted variables (Altonji, Elder, and Taber, 2005; Oster, 2019). In terms of economic effects, the 1-year probability of default of a loan increases by 16.1% in relative terms when interest rates were cut before loan origination (given that the average default probability equals 1.9 percentage points). Moreover, a 1 percentage point change in the interest rate after loan origination increases the probability of loan delinquency by 14.4% (following column 6).

In column 7 we test for the role of the U shape in monetary policy rates with the introduction of the interaction term between Cut and ΔRate. The estimated coefficient on this variable is positive and economically significant, consistent with our hypothesis. Periods of loose monetary policy followed by a 1 percentage point increase in the policy rate raise the probability of loan default by 28.4%. In columns 8 and 9 we replicate columns 6 and 7, but restrict the sample to those observations with observable firm financial characteristics. This leads to very similar results for the interaction estimated effects (though the interaction coefficient more than doubles the effect of just increasing monetary policy rates but without a previous cut in rates).

In Table 9 we explore the existence of heterogeneous effects of the U shape of monetary policy by interacting the term Cut × ΔRate with loan, firm and bank characteristics. From column 1 it follows that this effect is stronger for short-term loans (maturity shorter than 1 year). For these loans the relative increase in the likelihood of default reaches 44.5%. Columns 2 and 4 show that the effect of the U shape is more relevant for risky firms, i.e., those with a bad credit history or those that did not report to the Mercantile Register in the previous year (which means that the firm was not audited and, hence, had no administrative data for that year). For this group, the probability of default increases by an extra 55% if rates are raised by 1 percentage point after a period of cuts. For construction and real estate firms, this figure is somewhat higher at 60.5% (column 3). From the bank side, the observed U-shaped effect is stronger for riskier banks (those with higher non-performing loan ratios). For instance, going from the 25th to 75th percentile of the NPL ratio distribution increases the likelihood of default after U-shaped monetary policy by 82% (column 5).

Column 6 tests the robustness of all previous results by including all the interaction terms contemporaneously in the same regression. This leaves the results almost unchanged. Moreover, column 7 controls for time fixed effects, which absorbs all linear monetary policy effects and other non-observable time-varying macroeconomic variables such as realized and expected output growth, inflation, and financial risk. After controlling for these observable and unobservable variables, the heterogeneous effects are maintained. As in Table 8, the last two columns replicate columns 6 and 7 for the sample merged with the Mercantile Register database (which contains administrative data on firms' balance sheets). For this sample we have another proxy for the risk of a firm: its average cost of bank credit, with a higher cost of credit corresponding to, on average, riskier firms. Columns 8 and 9 show that the U shape in monetary policy affects these firms more. Going from the 25th to 75th percentile of the distribution of the average cost of credit raises the default probability

Table 9: *The path of monetary policy and loan-level defaults in Spain: heterogeneous effects*

	Dependent variable: Loan default at t+1								
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$\Delta\text{Rate}_{t,t+1}$	0.002*	0.002**	0.002**	0.002**	0.002*	0.002**		0.001*	
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)		(0.001)	
Cut $\text{Rate}_{t-5,t}$	0.003**	0.003**	0.003*	0.002	0.002	0.003**		0.004***	
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)		(0.001)	
$\Delta\text{Rate}_{t,t+1} \times$ Cut $\text{Rate}_{t-5,t}$	0.002*	0.002*	0.002*	0.002*	0.007***	0.005***		0.005***	
	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)		(0.001)	
ΔRate $\times$ Cut$\times$Short maturity	0.003***					0.004***	0.004***	0.003***	0.003***
	(0.001)					(0.001)	(0.001)	(0.001)	(0.001)
ΔRate $\times$ Cut$\times$Bad credit history		0.005**				0.005**	0.005**	0.003*	0.004*
		(0.002)				(0.002)	(0.002)	(0.002)	(0.002)
ΔRate $\times$ Cut$\times$Firm cost of credit								0.001***	0.001***
								(0.000)	(0.000)
ΔRate $\times$ Cut$\times$Real estate firm			0.006**			0.007***	0.007***	0.004*	0.004*
			(0.002)			(0.002)	(0.002)	(0.002)	(0.002)
ΔRate $\times$ Cut$\times$Not Mercantile Reg.				0.005***		0.005***	0.004**		
				(0.002)		(0.001)	(0.002)		
ΔRate $\times$ Cut$\times$Bank NPL Ratio					0.189***	0.180***	0.167***	0.136***	0.129***
					(0.047)	(0.043)	(0.056)	(0.036)	(0.040)
Bank Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm*Bank Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time Fixed Effects	No	No	No	No	No	No	Yes	No	Yes
Firm Controls	No	No	No	No	No	No	No	Yes	Yes
R^2	0.318	0.324	0.318	0.318	0.318	0.327	0.328	0.353	0.353
Observations	1.75m	1.75m	1.75m	1.75m	1.75m	1.75m	1.75m	1.27m	1.27m

Notes: This table reports the OLS regression results of the probability that a loan granted at time t becomes delinquent in the next year (t+1) for different paths of monetary policy rates. Cut is a dummy variable indicating whether the change in (overnight) rates between years t-5 and t is below its average value, and ΔRate is the percentage point change in the policy rate between years t and t+1. Coefficients are listed in the first row, robust standard errors are reported in the row below. Standard errors are corrected for clustering at the firm, bank and time level. Short maturity means loan maturity of 1 year or less. Firm cost of credit means measures the average cost of bank credit for the specific firm. Real estate firm includes both construction and real estate firms. Not Mercantile Reg. means that the firm did not report to the Mercantile Register in the previous year. For observations, m corresponds to millions. *, **, and *** indicate significance at the 0.1, 0.05, and 0.01 levels, respectively.

due to the U shape of monetary policy by 47.5%.

Taken together, the results in this section provide an important link not detectable at the aggregate level. Controlling for a large set of bank and firm characteristics, we find that a loan is more likely to default when originated at a time when monetary policy rates had decreased over a five-year window, and that this effect is stronger when interest rates start increasing. This effect is stronger for loans with short maturities, for risky firms and those in the real estate sector, and for weak banks, pointing towards a strong role of these characteristics in driving the build-up of financial vulnerabilities during times of monetary easing.

6. Conclusion

We analyze the link between monetary policy cycles and financial stability using macro data for 17 countries going back to 1870, and detailed micro data covering the post-1995 period in Spain. In the macro data, we find that pre-crisis monetary policy follows a U shape. We show that rates are, generally, cut 7 to 3 years before the crisis, and then increased in the run-up to the crisis. This U shape holds across different crisis definitions, short-term rate measures, and becomes more pronounced over time: we find that every single deep crisis after World War 2 was preceded by a U-shape in monetary policy rates. When it comes to predicting crises, we show that only those monetary policy rate hikes that were preceded by a series of cuts robustly increase crisis risk. The interest rate U shape is much more prominent for nominal short-term rates than for inflation, real rates, or long-term rate measures. It is also much more prominent before financial crises as opposed to non-crisis recessions.

To understand why U-shaped monetary policy is linked to crises, we first show that the initial loosening of policy is followed by high growth in credit and asset prices, putting the economy into a vulnerable financial "red zone" of Greenwood et al. (2022). After the subsequent monetary tightening, these "red-zone" vulnerabilities materialize, leading to larger-than-usual credit and asset price declines, higher crisis risk, and larger than usual declines in real activity. To dig further into the underlying mechanisms, we use administrative data on the universe of bank loans and defaults during the 1990s and 2000s boom-bust cycles in Spain. Consistently, we find that U-shaped monetary policy increases the probability of ex-post loan defaults, but effects are much stronger for ex-ante riskier firms (including real estate firms) and for banks with weaker balance sheets.

Overall, our analysis shows that the dynamic path of monetary policy has important implications for financial stability. Our paper does not offer a normative implication, but it suggests some important trade-offs, pointing towards a very subtle and nuanced view on the use of monetary policy to mitigate financial stability risks. For example, Greenwood et al. (2022) suggest that in order to reduce crisis risk, policymakers can raise interest rates to lean against the wind and lower credit and asset price growth. Our findings show that if monetary policy rates have been low for some time, allowing financial risks to build up, increasing policy rates would actually crystallize these vulnerabilities and dramatically increase crisis risk. Instead, deflating booms as early as possible and not allowing countries to enter into the "red zone" by having low monetary policy rates for long might be more appropriate.

But why then not focus on the financial vulnerabilities directly, instead of looking at the path of past monetary policy? One reason could be that monetary policy "gets in all the cracks" (Stein, 2013), and is able to affect financial risks in markets and segments not considered by regulators and macroprudential policy. For example, our results point to importance of monetary policy dynamics for both firms and households, and for both stock prices and house prices. The recent boom-and-bust cycle in crypto assets as well as many other financial markets (see The Economist, 8th December 2022) suggests that monetary policy affects seemingly distant financial markets where instabilities may build up outside of regulatory oversight.

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Appendix

A. Monetary policy and inflation around crises: Additional results

Residual policy rates We separate interest rate movements into endogenous and residual according to common specifications of the Taylor rule, and then plot the residual rate path around crises. We consider both the simple Taylor rule variant where policy responds to inflation and GDP, and a "kitchen sink" variant where we account for time-period breaks in monetary policy stance, and allow policy to respond to both financial (i.e., credit) and macro variables. Because we want to look at the entire rate path from $t-7$ to t, we do not condition on previous year's interest rate in the estimation. Our approach is therefore different to previous studies which used Taylor rules to separate yearly or quarterly changes in interest rates into systematic central bank responses and policy shocks consisting of random variation (Clarida, Galí, and Gertler, 2000).

Table A.1 shows the estimated Taylor rule coefficients for the full sample and post World War 2 period. In line with the literature, we find that interest rates correlate positively with inflation, and negative—with GDP growth, with these interest rate responses becoming stronger after 1975 (Sims and Zha, 2006). Finally, we add to these estimates a proxy for exogenous interest rate variation that is independent of any potential formulation of the Taylor rule, and is therefore robust to any rule misspecification errors. To do this, we take the exogenous interest rate movements in the Jordà et al. (2020) trilemma instrument in any given year, and cumulate them up from years $t-7$ to $t+7$ around the start of the crisis. These policy changes correspond to rate movements in pegged exchanged rate regimes with open capital markets, undertaken in order to maintain the peg (e.g., Spain adjusting in response to Germany under the ERM).

We then run the crisis window regression in equation (1) with these three definitions of residual rates as the dependent variable. The estimated regression coefficients for different pre- and post-crisis horizons, β_h, alongside the 90% confidence intervals, are shown in Figure A.1. Panel (a) uses the residual from the simple Taylor rule specification in column 1 of Table A.1, panel (b) uses the residual from the regression in column 2 of Table A.1, and panel (c) uses the cumulative changes in the Jordà et al. (2020) trilemma instrument. As in Figure 3, we show the coefficients for three different definitions of the crisis dummy: all systemic banking crises catalogued by Jordà et al. (2016) (left panel), deep Jordà et al. (2016) crises (middle panel), and post World War 2 crises (right panel). Across all the different crisis and residual rate definitions, policy rates follow the same U-shaped pattern as that catalogued in Figures 2 and 3a. Again, the residual rate variation is large, and both economically and statistically significant.

Long-term rates Figure A.2 shows the coefficients from regressing the long-term government bond yield (panel a) and the term premium (the long minus the short-term rate, panel b) on the crisis dummy for horizons $h = -7, ..., 7$ in the crisis window regression in (1). The long-term rate response is much weaker than that of the short-term rate in Figure 3a, and the term premium response shows a Λ rather than a U shape before crises.

Recessions Figure A.3 shows the recession window regressions for the path of policy rates, inflation, and real interest rates. These regression follow the crisis window set-up shown in equation (1), but use a recession (business cycle peak) dummy on the right-hand side instead of the crisis dummy. There is evidence that interest rates increase before a recession, but compared to crises the interest rate path is much more flat and there is little evidence of the U (i.e., there is no clear evidence of declining interest rates in years $t-7$ to $t-3$ before a recession).

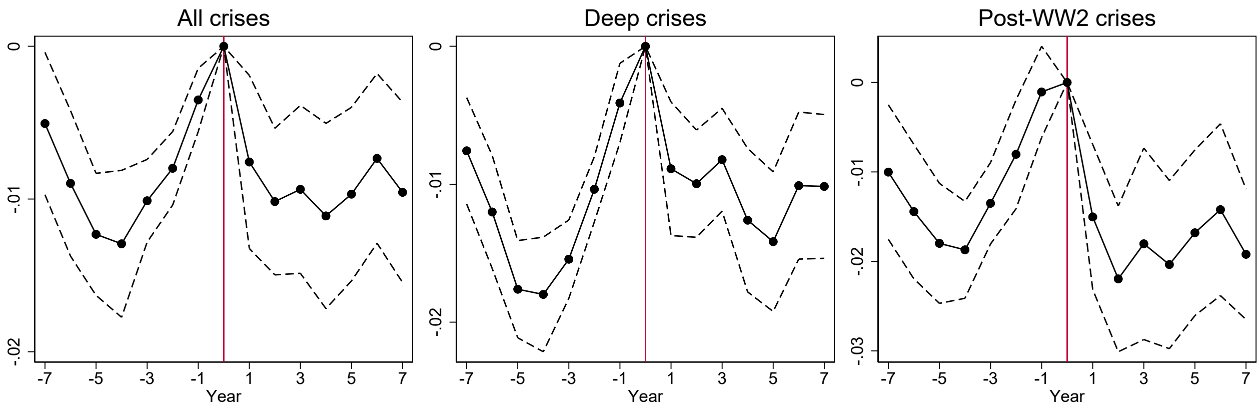
Table A.1: *Interest rates and macro-financial covariates*

	Dependent variable: policy rate			
	Full sample		Post-1945	
	(1)	(2)	(3)	(4)
Inflation	0.04***	0.04	0.19***	0.36***
	(0.01)	(0.03)	(0.03)	(0.07)
L. Inflation	0.03**	0.11***	0.10	0.22***
	(0.01)	(0.02)	(0.06)	(0.06)
L2. Inflation	0.06***	0.09**	0.06	-0.01
	(0.01)	(0.03)	(0.05)	(0.06)
ΔReal GDP	0.01	-0.01	-0.06	-0.03
	(0.01)	(0.02)	(0.05)	(0.06)
L. ΔReal GDP	0.03**	-0.01	-0.02	-0.07
	(0.01)	(0.03)	(0.05)	(0.11)
L2. ΔReal GDP	0.05***	0.01	0.03	0.06
	(0.01)	(0.02)	(0.05)	(0.04)
ΔReal credit		0.00		0.05
		(0.02)		(0.04)
L. ΔReal credit		0.03		-0.00
		(0.02)		(0.04)
L2. ΔReal credit		0.04**		-0.01
		(0.02)		(0.02)
ΔReal house price		0.01		0.00
		(0.01)		(0.02)
L. ΔReal house price		0.02*		0.01
		(0.01)		(0.02)
L2. ΔReal house price		0.02*		-0.00
		(0.01)		(0.03)
ΔReal stock price		0.00		0.01**
		(0.00)		(0.00)
L. ΔReal stock price		0.01		0.01**
		(0.00)		(0.00)
L2. ΔReal stock price		0.01**		0.02***
		(0.00)		(0.00)
R^2	0.09	0.19	0.21	0.36
Observations	2303	1732	1188	1070

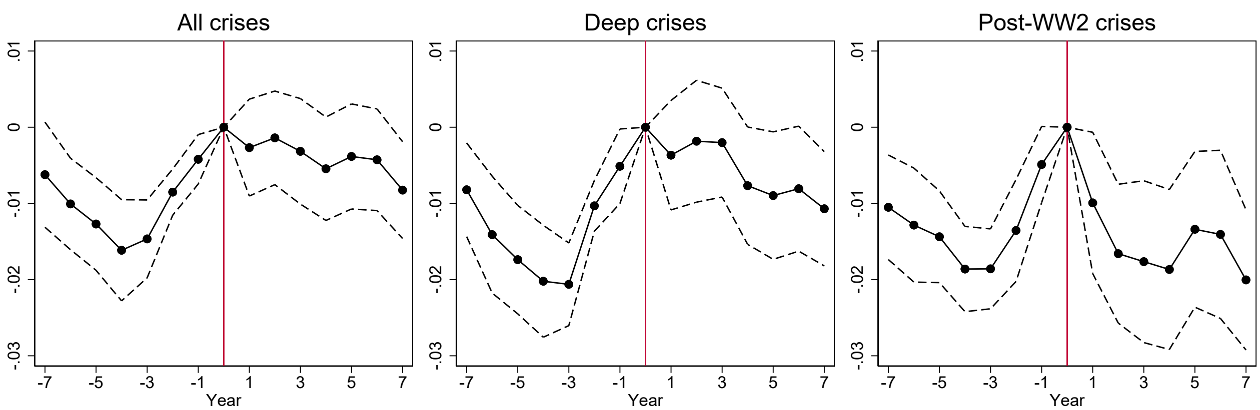
Notes: OLS regression with country fixed effects, 17 countries 1870–2016 (columns 1 and 2), and 1945–2016 (columns 3 and 4). All variables are for year *t*. Country-clustered standard errors in parentheses. *, **, and *** indicate significance at the 0.1, 0.05, and 0.01 levels, respectively.

Figure A.1: *Residual interest rates: Crisis window regressions*

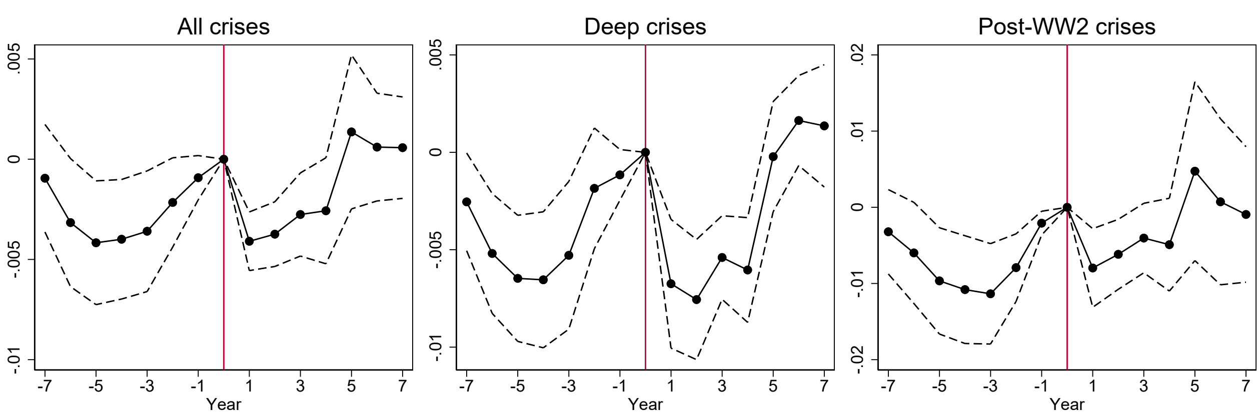
(a) *Simple*



(b) *Kitchen sink*



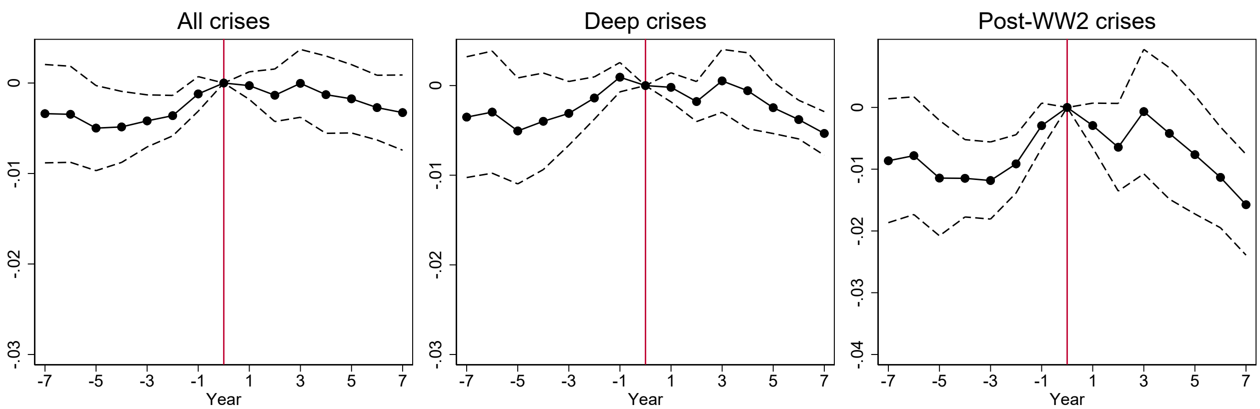
(c) *Cumulated trilemma IV*



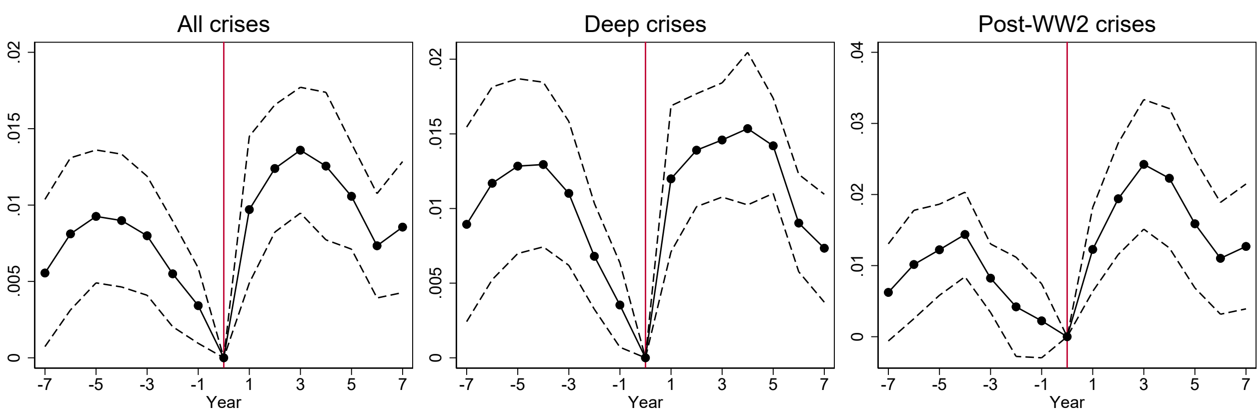
Notes: These graphs show the regression coefficients and 90% confidence intervals from regressing residual policy rates on the crisis dummy for horizons $h = -7, ..., 0, ...7$, with 0 corresponding to the beginning of the crisis according to the Jordà et al. (2016) chronology. Residual rates are measured as the residual from regressing the policy rates on contemporaneous values and two lags of inflation and GDP growth (panel a); the residual from regressing on inflation, GDP growth, credit growth, house price and stock price growth (again, contemporaneous and two lags, panel b); and by cumulating the exogenous changes in the Jordà et al. (2020) trilemma instrument (panel c). Deep crises are those with -3% or less GDP growth in one year, or average -1% or less GDP growth over 3 years in the $t-1$ to $t+3$ crisis window. Post-WW2 crises are those that started after 1945.

Figure A.2: *Long-term interest rates: Crisis window regressions*

(a) *Long-term interest rate*



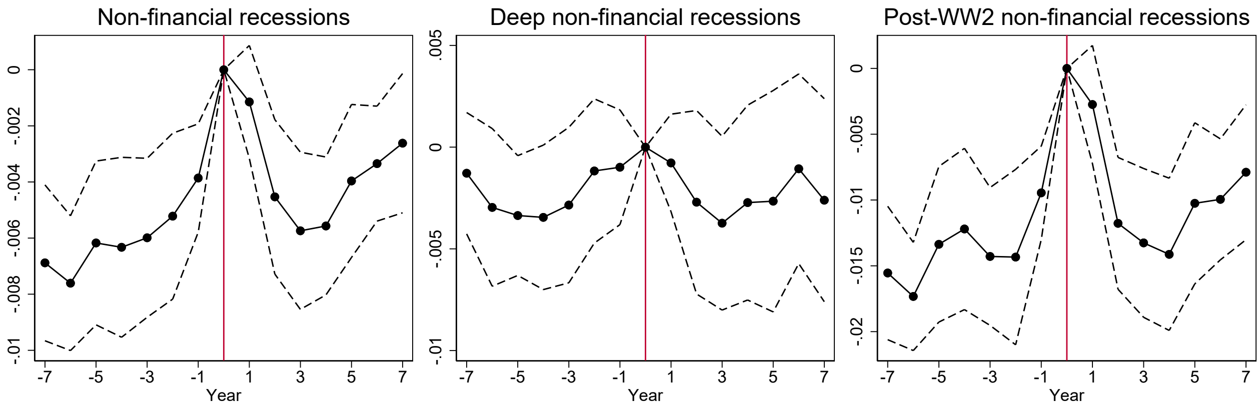
(b) *Term premium (long minus short rate)*



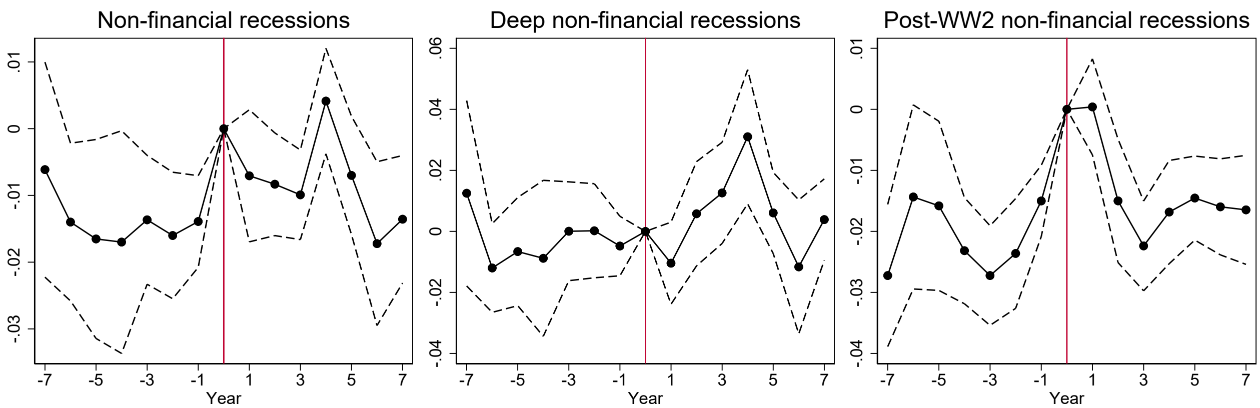
Notes: These graphs show the regression coefficients and 90% confidence intervals from regressing long-term rates and term premia on the crisis dummy for horizons $h = -7, ..., 0, ...7$, with 0 corresponding to the beginning of the crisis according to the Jordà et al. (2016) chronology. Long-term interest rate is the long-term government bond yield, targeting 10-year maturity. The term premium is the difference between long and short-term interest rates Deep crises are those with -3% or less GDP growth in one year, or average -1% or less GDP growth over 3 years in the $t - 1$ to $t + 3$ crisis window. Post-WW2 crises are those that started after 1945.

Figure A.3: *Policy rates, inflation, and real rates: Recession window regressions*

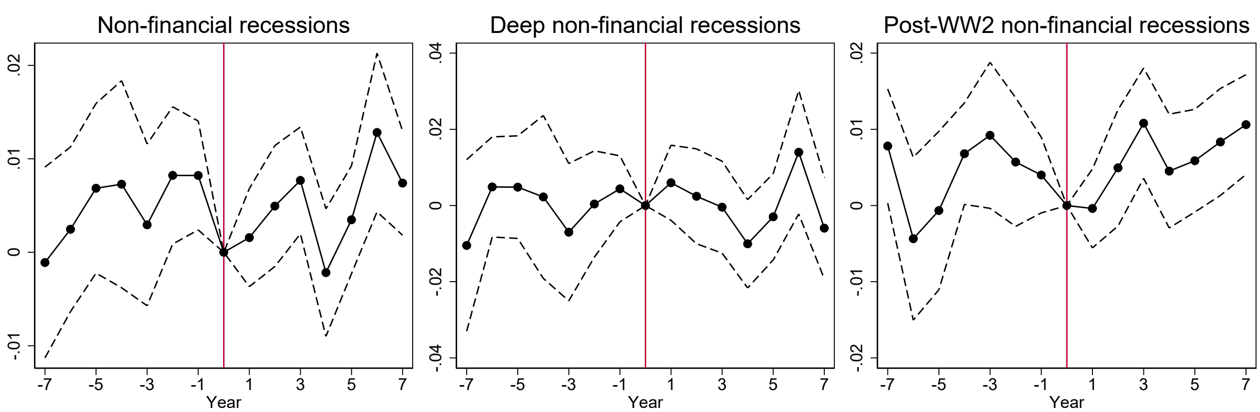
(a) *Policy rate*



(b) *Inflation*



(c) *Real interest rate*



Notes: These graphs show the regression coefficients and 90% confidence intervals from regressing residual policy rates on the non-crisis recession dummy for horizons $h = -7, ..., 0, ...7$, with 0 corresponding to the business cycle peak (identified using the Bry and Boschan, 1971 algorithm) that does not overlap with a systemic banking crisis dated according to the Jordà et al. (2016) chronology. Deep recessions are those with a -3% GDP or less growth in one year or average -1% or less growth over 3 years, in the 3-year window after the business cycle peak. Post-WW2 recessions are those that started after 1945.

B. Policy rates and crisis risk: Additional results

Table A.2: *Monetary policy shape and 1-year ahead crisis frequencies*

	(1)	(2)	(3)	(4)
	Crisis	Deep crisis	Post-WW2 crisis	Post-WW2 deep crisis
U shape (cut, raise)	0.07	0.04	0.06	0.05
Raise, raise	0.03	0.01	0.01	0.00
Raise, cut	0.02	0.01	0.00	0.00
Cut, cut	0.01	0.01	0.01	0.00
Unconditional	0.03	0.02	0.02	0.01

Notes: This table reports the crisis probability in year $t+1$ for different crisis definitions. In rows, the 4 bins are defined by the sign of the change (cut or raise) in the nominal policy rate between $t-8$ and $t-3$ and sign of the change (cut or raise) between $t-3$ and t. For example U-shape (cut, raise) refers to a cut in rates between $t-8$ and $t-3$ and a subsequent raise in rates between $t-3$ and t. Crisis frequency is the ratio of crisis to total observations in those years. Crises are dated using the Jordà et al. (2016) chronology. Deep crisis is crisis accompanied by at least -3% GDP growth in 1 year, or -1% average growth over 3 years in the window $t-1$ to $t+3$ around the crisis. Post-WW2 crises are those which started after 1945.

Table A.3: *Monetary policy shape and non-crisis recession frequencies*

	(1)	(2)	(3)	(4)
	Non-crisis recession	Deep non-crisis recession	Post-WW2 non-crisis recession	Post-WW2 deep non-crisis recession
U shape (cut, raise)	0.37	0.15	0.25	0.04
Raise, raise	0.30	0.12	0.27	0.05
Raise, cut	0.28	0.11	0.21	0.02
Cut, cut	0.26	0.15	0.09	0.00
Unconditional	0.31	0.13	0.21	0.03

Notes: This table reports the probability of a non-crisis recession (Bry and Boschan, 1971 business cycle peak that is not accompanied by a crisis) between year t and $t+2$ for different recession definitions and paths of nominal monetary policy rates. Recession frequency is the ratio of recession to total observations in those years. Deep recession is accompanied by at least -3% GDP growth in 1 year, or -1% average growth over 3 years in the window $t-1$ to $t+3$ around the business cycle peak. Post-WW2 recessions are those which started after 1945. In rows, the 4 bins are defined by the sign of the change (cut or raise) in the nominal policy rate between $t-8$ and $t-3$ and sign of the change (cut or raise) between $t-3$ and t. For example U-shape (cut, raise) refers to a cut in rates between $t-8$ and $t-3$ and a subsequent raise in rates between $t-3$ and t.

Table A.4: *Path of real interest rates and crisis frequencies*

	(1)	(2)	(3)	(4)
	Crisis	Deep crisis	Post-WW2 crisis	Post-WW2 deep crisis
U shape (fall, increase)	0.12	0.08	0.09	0.08
Increase, increase	0.09	0.02	0.07	0.03
Increase, fall	0.06	0.03	0.04	0.01
Fall, fall	0.07	0.02	0.06	0.01
Unconditional	0.09	0.04	0.06	0.03

Notes: This table reports the crisis probability between t and $t+3$ for different crisis definitions, for different paths of the real interest rate. In rows, the 4 bins are defined by the sign of the change (fall or increase) in the real interest rate between $t-8$ and $t-3$, and sign of the change (fall or increase) between $t-3$ and t. For example U-shape (fall, increase) refers to a fall in real rates between $t-8$ and $t-3$ and a subsequent increase between $t-3$ and t. Crisis frequency is the ratio of crisis to total observations in those years. Crises are dated using the Jordà et al. (2016) chronology. Deep crisis is crisis accompanied by at least -3% GDP growth in 1 year, or -1% average growth over 3 years in the window $t-1$ to $t+3$ around the crisis. Post-WW2 crises are those which started after 1945.

Table A.5: *Path of inflation and crisis frequencies*

	(1)	(2)	(3)	(4)
	Crisis	Deep crisis	Post-WW2 crisis	Post-WW2 deep crisis
U shape (fall, increase)	0.09	0.06	0.09	0.06
Increase, increase	0.07	0.03	0.06	0.03
Increase, fall	0.10	0.06	0.04	0.03
Fall, fall	0.05	0.01	0.01	0.00
Unconditional	0.08	0.05	0.05	0.03

Notes: This table reports the crisis probability between t and $t+3$ for different crisis definitions, for different paths of inflation. In rows, the 4 bins are defined by the sign of the change (fall or increase) in inflation between $t-8$ and $t-3$, and sign of the change (fall or increase) between $t-3$ and t. For example U-shape (fall, increase) refers to a fall in inflation between $t-8$ and $t-3$ and a subsequent increase between $t-3$ and t. Crisis frequency is the ratio of crisis to total observations in those years. Crises are dated using the Jordà et al. (2016) chronology. Deep crisis is crisis accompanied by at least -3% GDP growth in 1 year, or -1% average growth over 3 years in the window $t-1$ to $t+3$ around the crisis. Post-WW2 crises are those which started after 1945.

Table A.6: *Crisis prediction models: Robustness (IV)*

	1-year ahead		Post-WW2		Driscoll-Kraay		Decade FE + Controls	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta_3\text{Rate}_t$	0.01*	0.00	0.03***	0.00	0.03	0.00	0.02	0.00
	(0.01)	(0.00)	(0.01)	(0.01)	(0.02)	(0.01)	(0.02)	(0.02)
Cut $\text{Rate}_{t-8,t-3}$	0.02**	0.02**	0.06***	0.06***	0.06*	0.06*	0.03	0.04
	(0.01)	(0.01)	(0.02)	(0.02)	(0.03)	(0.03)	(0.03)	(0.03)
$\Delta_3\text{Rate}_t$ × Cut $\text{Rate}_{t-8,t-3}$		0.03**		0.09***		0.07*		0.06***
		(0.01)		(0.03)		(0.04)		(0.02)
Country fixed effects	✓	✓	✓	✓	✓	✓	✓	✓
Kleibergen-Paap Weak ID	95.95	42.94	86.12	54.64	48.29	24.17	48.21	17.65
Observations	1625	1625	949	949	1625	1625	1198	1198

Notes: This table shows linear probability models for a systemic banking crisis occurring in year $t+1$ (columns 1 and 2), or years t to $t+2$ (columns 3–8). Post-WW2 specifications restrict the sample to after 1945. Driscoll-Kraay specifications cluster standard errors at country and year level, allowing for autocorrelation of up to 5 lags. Decade FE + Controls specifications include both country and decade fixed effects, and control for eight lags of real credit, house price, and stock price growth, as well as inflation and real GDP growth. Specifications in columns 1–6 control for 8 lags of GDP growth and inflation. All specifications instrument policy rate changes with the residualized Jordà et al. (2020) trilemma variable. Interactions are between policy rate changes and a dummy for a cut between $t-8$ and $t-3$. IV interaction specifications include residualized JST trilemma variable and its interaction with the cut dummy as instruments. In this case the KP Weak ID is the joint test for both instruments. Standard errors in columns 1–4 and 7–8 are clustered by country and year. *, **, and *** indicate significance at the 0.1, 0.05, and 0.01 levels, respectively.

Table A.7: *Crisis prediction: probit*

	OLS			IV		
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta_3\text{Rate}_t$	0.13***	0.12***	0.06**	0.23***	0.21**	0.01
	(0.03)	(0.03)	(0.03)	(0.08)	(0.08)	(0.10)
Cut $\text{Rate}_{t-8,t-3}$		0.40**	0.36**		0.33*	0.34**
		(0.16)	(0.16)		(0.17)	(0.17)
$\Delta_3\text{Rate}_t$ × Cut $\text{Rate}_{t-8,t-3}$			0.13***			0.36***
			(0.03)			(0.13)
Country fixed effects	✓	✓	✓	✓	✓	✓
Observations	1564	1564	1564	1564	1564	1564

Notes: This table shows probit models for a systemic banking crisis occurring in years t to $t+2$. The IV specifications instrument policy rate changes with the residualized Jordà et al. (2020) trilemma variable. Interactions are between policy rate changes and a dummy for a cut between $t-8$ and $t-3$. IV interaction specifications include residualized JST trilemma variable and its interaction with the cut dummy as instruments. Country-clustered standard errors in parentheses. All specifications control for 8 lags of GDP growth and inflation. *, **, and *** indicate significance at the 0.1, 0.05, and 0.01 levels, respectively.

Table A.8: *Crisis prediction: low-for-long interest rates*

	Crisis in t+1 to t+3			
	OLS		*IV*	
	(1)	(2)	(3)	(4)
ΔRate_t	0.02***	0.01*	0.03**	-0.00
	(0.00)	(0.00)	(0.01)	(0.01)
No. years (low spell)$_{t-1}$	0.00	0.00	0.01**	0.00
	(0.00)	(0.00)	(0.00)	(0.00)
$\Delta \text{Rate}_t \times$ No. years (low spell)$_{t-1}$		0.01***		0.02***
		(0.00)		(0.01)
Country fixed effects	✓	✓	✓	✓
Controls	✓	✓	✓	✓
Kleibergen-Paap Weak ID			55.84	25.11
Observations	1865	1865	1614	1614

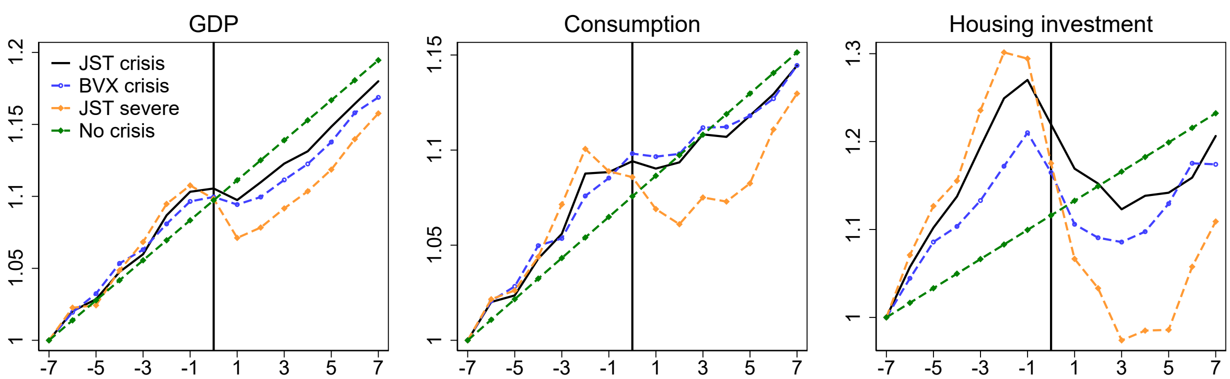
Notes: ΔRate_t is the yearly change in monetary policy rates. No. years low spell measures the number of years policy rates have been lower than their average over a preceding 10-year window. IV specifications instrument ΔRate_t and its interaction with the number of years in a low spell with the Jordà et al. (2020) trilemma instrument and its interaction with the low spell. All specifications control for 8 lags of GDP growth and inflation. Country-clustered standard errors in parentheses. *, **, and *** indicate significance at the 0.1, 0.05, and 0.01 levels, respectively.

C. Mechanisms in macro data: Additional results

Figure A.4: *Credit, asset prices, and real activity around different types of crises*

(a) *Credit and asset prices*

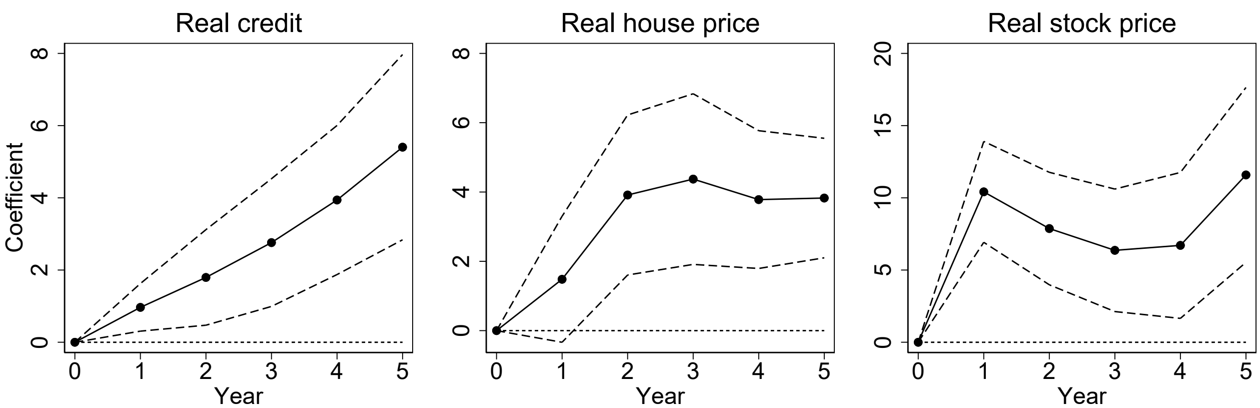
(b) *Real activity*



Notes: Unweighted averages of 72 systemic banking crises defined as in Jordà et al. (2016). To construct the figures, we calculate average log growth rates of each variable in a given pre- or post-crisis year, and cumulate them starting at 1, from 7 years before to 7 years after the crisis.

Figure A.5: *Response of credit and asset prices to policy rate cuts, alternative specifications (IV)*

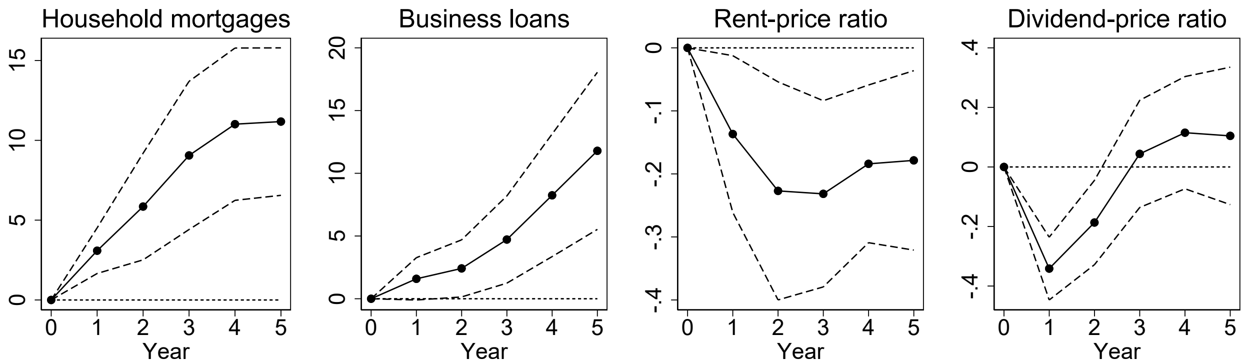
(a) *Post World War 2*



(b) *Large cuts in policy rates*

Notes: Local projection estimates of future credit, house price, and stock price growth, for horizons $t+1$ to $t+h$ on a cut in interest rates at t. Large cuts consist of all changes in interest rates (predicted using the Jordà et al. (2020) trilemma instrument) that are in the lowest quartile of the rate change distribution. All regressions include country and decade fixed effects, and control for contemporaneous levels and four lags of real credit, house price, stock price, GDP, consumption, and housing investment growth, as well as two four of changes in short- and long-term rates. Dashed lines show 90% confidence intervals.

Figure A.6: *Response of household and business lending and risk premia to policy rate cuts (IV)*



Notes: Local projection estimates of future real loan growth and risk premia, for horizons $t+1$ to $t+h$ on interest rate changes at t normalized to a 1 percentage point cut. All policy rate changes are instrumented using the Jordà et al. (2020) trilemma instrument. All regressions include country and decade fixed effects, and control for contemporaneous levels and four lags of real credit, house price, stock price, GDP, consumption, and housing investment growth, as well as two four of changes in short- and long-term rates. Dashed lines show 90% confidence intervals.

Table A.9: *R-zone and crisis risk: validation in our sample*

	Dependent variable: $\text{Crisis}_{t\ to\ t+2}$		
	(1)	(2)	(3)
R-zone $\text{Business}_{i,t}$	0.12^{**}		
	(0.06)		
R-zone $\text{Households}_{i,t}$		0.21^{***}	
		(0.05)	
R-zone $\text{Either}_{i,t}$			0.14^{***}
			(0.05)
Country fixed effects	✓	✓	✓
Controls	✓	✓	✓
Observations	1813	1578	1875

Notes: R-zone is defined as joint high growth in credit and asset prices, using the same definition as in Greenwood et al. (2022). We use business credit and equity prices for the business R-zone, and household credit and house prices for the household R-zone. Country-clustered standard errors in parentheses. *, **, and *** indicate significance at the 0.1, 0.05, and 0.01 levels, respectively.

Table A.10: *Policy rate reversal, credit, and asset prices: OLS*

	Loans		House Prices		Stock Prices	
	3-yr	6-yr	3-yr	6-yr	3-yr	6-yr
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta_3 \text{Rate}_t$	-1.04***	-1.66***	-1.21***	-1.25***	-2.72***	-1.44
	(0.19)	(0.44)	(0.28)	(0.42)	(0.78)	(1.15)
Cut $\text{Rate}_{t-8,t-3}$	-0.02	-0.03	0.01	0.02	-0.03	-0.02
	(0.01)	(0.02)	(0.01)	(0.02)	(0.03)	(0.03)
$\Delta_3 \text{Rate}_t \times$ Cut $\text{Rate}_{t-8,t-3}$	-0.96***	-1.32***	-0.63**	-1.54***	-0.79	0.96
	(0.20)	(0.35)	(0.27)	(0.44)	(0.57)	(1.10)
Observations	1054	1054	1061	1061	1056	1060

Notes: y variables in columns refer to growth in the real variable (e.g., loans/CPI) between $t-1$ and $t+2$ (3-yr specifications), and $t-1$ and $t+5$ (6-yr specifications). x variable is the raw change in the policy rate. All regressions include country and decade fixed effects, and control for four lags of real loan, house price, stock price, GDP, consumption, and housing investment growth, and four lags of the crisis dummy. Country-clustered standard errors in parentheses. *, **, and *** indicate significance at the 0.1, 0.05, and 0.01 levels, respectively.

Table A.11: *Policy rate reversal and real macroeconomic outcomes: OLS*

	GDP		Consumption		Housing investment	
	3-yr	6-yr	3-yr	6-yr	3-yr	6-yr
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta_3 \text{Rate}_t$	-0.41***	-0.46***	-0.47***	-0.52***	-2.20***	-1.99***
	(0.08)	(0.18)	(0.08)	(0.20)	(0.42)	(0.55)
Cut $\text{Rate}_{t-8,t-3}$	-0.00	0.01	0.00	0.01	-0.01	-0.00
	(0.00)	(0.01)	(0.01)	(0.01)	(0.03)	(0.02)
$\Delta_3 \text{Rate}_t \times$ Cut $\text{Rate}_{t-8,t-3}$	-0.33***	-0.40**	-0.13	-0.09	-1.12*	-1.54**
	(0.12)	(0.18)	(0.13)	(0.17)	(0.60)	(0.74)
Observations	1066	1066	1066	1066	1045	1045

Notes: y variables in columns refer to growth in the real variable (e.g., real GDP) between $t-1$ and $t+2$ (3-yr specifications), and $t-1$ and $t+5$ (6-yr specifications). x variable is the raw change in the policy rate. All regressions include country and decade fixed effects, and control for four lags of real loan, house price, stock price, GDP, consumption, and housing investment growth, and four lags of the crisis dummy. Country-clustered standard errors in parentheses. *, **, and *** indicate significance at the 0.1, 0.05, and 0.01 levels, respectively.

D. Loan-level evidence from Spain: Additional results

Table A.12: *Spanish administrative data: summary statistics*

		Mean (1)	S.D. (2)	P25 (3)	Median (4)	P75 (5)
Loan default$_{t,t+1}$	0/1	0.019	0.135	0.000	0.000	0.000
ΔRate$_{t,t+1}$	%	-0.326	1.093	-0.906	-0.143	0.245
Cut Rate$_{t-5,t}$	0/1	0.427	0.495	0.000	0.000	1.000
Short maturity	0/1	0.503	0.500	0.000	1.000	1.000
Firm bad credit history	0/1	0.109	0.311	0.000	0.000	0.000
Construction and real estate firm	0/1	0.214	0.410	0.000	0.000	0.000
Firm not in Mercantile Register the previous year	0/1	0.246	0.431	0.000	0.000	0.000
Firm average cost of credit	%	3.190	2.801	1.052	2.597	4.610
Bank NPL Ratio	0.0x	0.043	0.051	0.008	0.017	0.061

Notes: This table reports means, standard deviations and first/second/third quartiles of the main variables used in the loan-level regressions using Spanish administrative data.