Loose Monetary Policy and Financial Instability*

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Abstract

Do periods of persistently loose monetary policy increase financial fragility and the likelihood of a financial crisis? This is a central question for policymakers, yet the literature does not provide systematic empirical evidence about this link at the aggregate level. In this paper we fill this gap by analyzing long-run historical data. We find that when the stance of monetary policy is accommodative over an extended period, the likelihood of financial turmoil down the road increases considerably. We investigate the causal pathways that lead to this result and argue that credit creation and asset price overheating are important intermediating channels.

JEL classification: E43, E44, E52, E58, G01, G21, N10

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1 Introduction

Does persistently loose monetary policy breed financial fragility? And if so, why? Scholars and policymakers alike blamed loose monetary policy for the boom-bust that culminated in the Global Financial Crisis (Geithner, 2009; Taylor, 2011) and warned yet again in its aftermath "that a long period of low interest rates... could undermine financial stability" (Bernanke 2013; also, see Stein 2013). However, despite the large adverse macroeconomic (Cerra and Saxena, 2008; Reinhart and Rogoff, 2009; Jordà, Schularick, and Taylor, 2013) and political (Funke, Schularick, and Trebesch, 2016; Doerr, Gissler, Peydró, and Voth, 2022) consequences of financial crises, there is no systematic empirical study that analyzes the link between the stance of monetary policy and macro-level financial stability.

This is not to say that the question has been completely ignored by the literature. On the contrary, there have been many micro-level empirical studies (to be discussed shortly) showing the causal effect of loose monetary policy on increased risk-taking by financial institutions and households. Such behavior by individual financial market participants can be theoretically justified (as we note below). However, it is unclear how individual actions aggregate up. We still do not know if this well-identified, micro-level evidence translates into measurable, or even dangerous, macro-level financial instability (Boyarchenko, Favara, and Schularick, 2022). This study fills this gap in the literature using macro-financial data for advanced economies over the past 150 years.

How do we know if interest rates are "too low for too long"—or, more specifically, if monetary policy is too loose? Ever since Wicksell (1898), macroeconomists have generally understood the equilibrium or natural real rate r^* to be that which leaves a fully flexible economy at full employment with stable inflation. Thus, as in the literature on policy rules, deviations of the real policy rate from the *natural rate* are a measure of monetary policy stance, or $stance = r - r^*$. A loose monetary policy is when $stance < 0$. Thus, to operationalize this idea, an important element of our study is to put together, as a first step, measures of r^* for the very long run.

Building on that, we are the first to show that, as a causal matter, a loose stance has strong implications for medium-term financial instability. Using an instrumental variable approach, when stance is 1 percentage point (pp) lower on average in a 5-year window, then the probability of a financial crisis in the next 5 to 7 years increases by 5.5 pps, and by 15.5 pps in the following 7 to 9 years ahead. Since the unconditional probability of experiencing a crisis in any 3-year window is 10.5 percent, these effects are big. Moreover, these results are robust to alternative measures of stance and alternative definitions of financial stability.

In the second part of the paper, we then go one step back and ask: *how* does excessively loose monetary policy trigger financial instability? Here we explore the likely mechanisms at work, looking into the suspected channels through which an accommodative stance translates into increased financial fragility, focusing on the risks of overheating in credit and asset markets.

Long ago, Kindleberger (1978) noted that "Speculative manias gather speed through expansion of money and credit or perhaps, in some cases, get started because of an initial expansion of money and credit" (p. 52). Why, though, do money and credit expand in the first place? By analyzing this question, we contribute to the strand of the literature that focuses on potential causes of credit and asset price booms. To the best of our knowledge, this strand is relatively thin.[1] However, the question has been in scholars' minds for a long time. The originator of the natural rate concept, Wicksell (1898), hypothesized that low interest rates—and low-for-long periods in particular—spur house prices (p. 88). He even went further and argued that such increases in house prices could generate feedback as entrepreneurs expect further price increases (p. 88). Eventually, speculation starts to dominate markets (pp. 89–90), resulting in a boom-and-bust cycle (p. 90). Such a mechanism running from low interest rates set by the central bank, through behavioral responses in credit quantities and asset prices, also figures in the recent model of Kashyap and Stein (2023).[2]

Consistent with these mechanisms, we provide evidence that overheated credit and asset markets can originate from accommodative monetary policy regimes. In line with the existing macro-finance literature, this then helps predict financial crises. Here, using the "Red-zone" definitions in Greenwood, Hanson, Shleifer, and Sørensen (2022), we find that when interest rates remain below the natural rate for an extended period of time, there is a buildup in asset prices and in credit growth, both of which have been shown to be associated with greater financial fragility (see, e.g., Schularick and Taylor, 2012; Jordà, Schularick, and Taylor, 2015a,b, 2016; Greenwood, Hanson, Shleifer, and Sørensen, 2022).

Most importantly, throughout the paper we aim to pin down the causal link between policy stance and financial fragility. Causal statements require a source of exogenous variation for endogenous variables, in this case the policy stance. Our analysis relies on an instrumental variable (IV) for *stance* based on the trilemma of international finance (Obstfeld

[1]See Mian and Sufi (2018, pp. 50–52) for an overview. Jordà, Schularick, and Taylor (2015a) and Jiménez, Kuvshinov, Peydró, and Richter (2022) provide evidence that a loosening of monetary conditions leads to rises in credit and asset prices. Bianchi, Lettau, and Ludvigson (2022) show in the context of the United States that long-lasting shifts in asset valuations coincide with shifts in the Fed's conduct of monetary policy.

[2]Kashyap and Stein (2023) write in this context: "as noted by Boyarchenko, Favara and Schularick (2022), there is limited evidence that it is specifically monetary-policy induced changes in credit growth and risk premiums—as opposed to changes driven by other factors—that create this economic vulnerability. As they note, establishing such a link is challenging, and more research would be welcome. We are going to make the leap and assume that the link is operative in what follows, but the reader should be aware that this connection is not yet firmly established." It is exactly this hypothesized link that is the object of study in this paper.

and Taylor, 2004) and used to study the effects of monetary policy by, e.g., Jordà, Schularick, and Taylor (2020). Here, as in Jordà, Schularick, and Taylor (2020), the differences between the OLS and IV results are stark, regardless of how we measure financial fragility, and a causal interpretation is justified.

In the last part of this paper, we turn attention to real activity outcomes. There, we show that too accommodative a stance of monetary policy is associated with a higher likelihood of left-tail "disaster" outcomes in the medium-term GDP growth distribution. This finding indirectly confirms the existing evidence already mentioned above that financial busts bear large macroeconomic costs. It also speaks to the literature that studies the impact of loose financial conditions on growth. For instance, Mian, Sufi, and Verner (2017) provide evidence that household debt booms are accompanied by a temporary boost in real activity. This boost, though, is short-lived and eventually reverses. Loose financial conditions are positive for the left tail of the predicted real GDP growth distribution in the short term, but at the expense of strong negative tail effects in the medium term without affecting the economy's expected growth path (Adrian, Boyarchenko, and Giannone, 2019; Adrian, Grinberg, Liang, Malik, and Yu, 2022). This suggests a trade-off between real activity and financial stability. Monetary policymakers face such a growth-risk trade-off in periods of low interest rates in the models of Acharya and Plantin (2018) and Coimbra and Rey (2023).

Theory and evidence at the micro level: risk taking and reach for yield Risk taking is likely to be affected by the stance of monetary policy (Bauer, Bernanke, and Milstein, 2023). As already noted earlier, several empirical studies show that when interest rates are low, financial institutions make riskier investment decisions, see, e.g., Maddaloni and Peydró (2011), Jiménez, Ongena, Peydró, and Saurina (2014), Altunbas, Gambacorta, and Marques-Ibanez (2014), Ioannidou, Ongena, and Peydró (2015), Dell'Ariccia, Laeven, and Suarez (2017), and Paligorova and Santos (2017) for banks. A similar link has been identified for other types of investors such as mutual funds (Choi and Kronlund, 2018), money market funds (Chodorow-Reich, 2014; Di Maggio and Kacperczyk, 2017), pension funds (Chodorow-Reich, 2014), and retail investors (Hau and Lai, 2016).

When interest rates are relatively loose, financial intermediaries have incentives—or are even required—to search for yield and thus risk. This incentive to "search for yield" was famously put forward by Rajan (2005) as one source of financial risk. One example he gave was insurance companies. These institutions often face fixed long-term commitments and therefore increase their risk appetite when rates are low. Hanson and Stein (2015) argue that there exist yield-oriented investors who not only care about the discounted return of their portfolio but place particular emphasis on current returns. Such investors increase their portfolio's maturity risk when short-term rates fall and vice versa. The authors find empirical support for their hypothesis: after expansionary monetary policy shocks, the maturity of banks' security holdings rises. Search-for-yield behavior was more recently rationalized by Campbell and Sigalov (2022). They add a sustainable spending constraint, which requires investors to consume the expected real return on their portfolio each period, to an otherwise standard model. The authors motivate this constraint with the behavior of sovereign wealth funds. It transforms the standard intertemporal trade-off between higher consumption today and *lower* future consumption into a trade-off between higher consumption today and *riskier* future consumption.

Drechsler, Savov, and Schnabl (2018) also establish a theoretical link between lower interest rates and increasing leverage and thus risk exposure. In their model, leveraged banks hold liquid securities to avoid the necessity to fire-sale assets when faced with an adverse funding shock. Hence, liquid securities trade at a premium that varies positively with the policy rate. Consequently, the opportunity cost of taking on leverage is lower during periods of low policy rates. This leads banks to increase risk taking in such environments. Where Drechsler, Savov, and Schnabl (2018) analyze liquidity holdings, Dell'Ariccia, Laeven, and Marquez (2014) endogenize capital structures, and capital serves as a signaling device for investors to gauge the degree of a bank's risk-taking behavior, thereby determining the external funding cost of a bank. Low interest rates reduce this benefit of capital as a signaling device. Low rates, therefore, lead to higher leverage ratios which in turn implies elevated risk taking, ceteris paribus. In their model, there is not only the aforementioned effect that works through banks' endogenous capital structure. The model also endogenizes banks' monitoring efforts, which it has in common with Martinez-Miera and Repullo (2017) and Heider and Leonello (2021). In all these studies, banks have fewer incentives to monitor borrowers when interest rates are low, implying higher risk taking. Expansionary monetary policy can also increase aggregate risk taking by redistributing wealth from conservative, risk-averse households toward leveraged households with higher risk tolerance, which, in turn, compresses risk premia (Kekre and Lenel, 2022).

Finally, from the experimental literature, Lian, Ma, and Wang (2019) find evidence for *reference dependence* and *salience*. In their experiments, an individual starting the experiment in a high interest rate environment will tend to make riskier investment decisions when shifted to a low interest rate environment. Through this lens, one could view the slow-moving trend level of the natural rate of interest r^* as a salient, headline, history-dependent reference level of the real rate, and deviations from it as the driver of risk taking.

To the best of our knowledge, Altunbas, Gambacorta, and Marques-Ibanez (2014) is the only study that attempts to separate monetary policy from secular trends in interest rates. The authors decompose real rates into trend and cycle components (with the HP Filter) and treat the trend as the natural rate, and the cycle as the stance of monetary policy. Instead, we will use a more fundamental measure of the natural rate based on Del Negro, Giannone, Giannoni, and Tambalotti (2019) from which to calculate the stance of monetary policy.

Mechanisms at the macro-level: aggregate consequences of a loose stance We now turn to the question as to why and how a loose stance can compound over time and across the economy as a whole so as amplify crisis risk.

Obviously, many of the channels explored in the micro literature could get stronger the longer that monetary policy remains loose, and several macro-level papers build on that idea. The danger of lower-for-longer monetary policy is stressed in Boissay, Collard, Galí, and Manea (2022). In their model, financial crises are the consequence of a central bank that keeps the policy rate too low for too long which in turn fosters an investment boom and eventually a capital overhang. Given this concern, we explicitly consider the consequences of *persistently* loose monetary policy as opposed to single periods of policy undershooting relative to the natural rate of interest. In Akinci, Benigno, and Del Negro (2020), persistent reductions in real interest rates lead to improved financial conditions and higher risk taking in the short term at the expense of increased financial fragility further down the road. On the empirical side, Jiménez, Kuvshinov, Peydró, and Richter (2022) find that financial crises are preceded by U-shaped monetary policy; policy rates characterized by cuts in the short term and hikes in the medium term are associated with higher crisis risk. Where they consider the nominal rate path, we analyze the stance of monetary policy, that is deviations of the real rate from the natural rate and how they evolve.

Our paper is also related to the literature on the classical balance sheet channel. Lower interest rates in general and loose monetary policy in particular imply, ceteris paribus, higher asset valuations (Bianchi, Lettau, and Ludvigson, 2022). This opens the door for collateral-driven credit booms (Kiyotaki and Moore, 1997). Adrian and Shin (2008) find empirical evidence for pro-cyclical bank leverage and for a positive effect of loose monetary policy on banks' balance sheet size. Based on these empirical findings, the authors then argue that the policy rate "may be a key price variable in its own right" that triggers a feedback effect between bank balance sheet strength, leverage, and asset prices, thereby creating an asset price boom. If a loose stance of monetary policy creates abundant liquidity, it can also contribute to the emergence of credit booms and asset price bubbles if bankers, motivated by volume-based compensation, take on excessive risks, ultimately "sowing seeds of the next crisis" (Acharya and Naqvi, 2012).

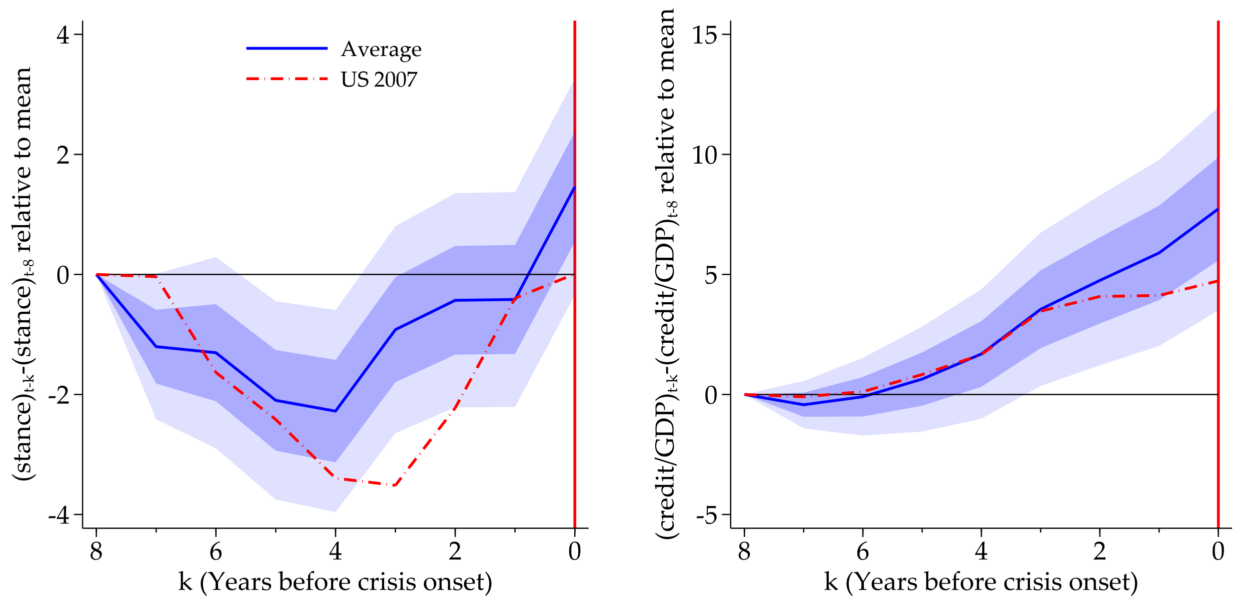
Indeed, such credit and asset price booms have been identified by the literature as harbingers of financial turmoil. Elevated credit growth (Schularick and Taylor, 2012; Jordà, Schularick, and Taylor, 2016), house price booms (Jordà, Schularick, and Taylor, 2015a), and their interaction (Jordà, Schularick, and Taylor, 2015b; Greenwood, Hanson, Shleifer, and Sørensen, 2022) are important short-run predictors of financial crises. Similarly, credit expansions predict bank equity crashes (Baron and Xiong, 2017).

Why is the stance of monetary policy loose in some periods and tight during others? The literature does not provide a definite answer to this question, and there may be a wide range of potential explanations. Consider, for instance, the stance of U.S. monetary policy in the 1970s which can be classified as loose (Clarida, Galí, and Gertler, 2000). Fed Chairman Burns' motivations for maintaining a loose stance of monetary policy were "complex," ranging from political pressure over systematic underestimation of the natural rate of unemployment to Burns' personal view on the causes of inflation (Burns, 1979; Bernanke, 2022, p. 29). Political pressure is also an explanation for the exceptionally loose stance of monetary policy of the Bank of England that preceded the first post-WWII financial crisis among major advanced economies (Reinhart and Rogoff, 2009; Jordà, Schularick, and Taylor, 2017), namely the British Secondary Banking Crisis of 1973–75 (Needham, 2014). Both in the U.K. and the U.S., these periods of accommodative monetary policy were accompanied by high inflation. However, a loose stance of monetary policy is not necessarily the consequence of monetary policymakers' deviation from the objective of price stability. A case in point is the most recent financial crisis in our sample. The Global Financial Crisis was also preceded by accommodative monetary policy, but not by excessive inflation rates. One explanation put forward for the loose stance of monetary policy during the 2000s is asymmetric behavior of central banks that react more strongly to financial busts than to booms (Hofmann and Bogdanova, 2012). Asymmetric real performance within a currency union is another source that can lead to deviations of monetary policy from its neutral rate. We take a closer look at pre-2008 monetary policy in the eurozone in section 4.

Motivation from an event study of stance, credit, and crises Using an event study approach and our measure of monetary policy stance (which we document in the next section) we can show preliminary evidence to motivate our work. Panel (a) of Figure 1 shows the path of *stance* as defined above before financial crises relative to normal times. The stance measure falls in the medium term and rises again in the immediate years before crisis events. For example, stance is on average 3.5 pps lower 5 years before a crisis compared to at the time of the crisis itself. Panel (b) illustrates the well-established evidence outlined above that credit booms are a precursor of financial busts. In the 5 years before a

Figure 1: *The stance of monetary policy and credit growth before financial crisis events.*

(a) *stance (pps)* (b) *Credit-to-GDP ratio (pps)*



Notes: In this figure, the data, including crisis event definitions, are taken from the JST Macrohistory Database, as described later. The solid blue line shows estimates of β_k of $\left(y_{i,t-k} - y_{i,t-8}\right) = \alpha_k + \beta_k \mathbb{1}\{crisis_{i,t} = 1\} + e_{t-k}$. $crisis_{i,t}$ is a dummy that is equal to 1 if a financial crisis starts in country i in year t and 0 otherwise. y refers to $stance = r - r^*$ (left panel), as defined in the text; or credit-to-GDP ratio (right panel), based on the JST total loans series. The estimation of r^* is described below in section 2. Shaded areas indicate 95% (light) and 68% (dark) confidence intervals. The dashed red line shows demeaned changes in the two variables before the U.S. Great Recession.

crisis event, the rise in the credit-to-GDP ratio is on average 1.7 pps per year higher than in normal times. For example, loose monetary policy characterized the macro-financial environment of the early 2000s in the United States, shown by the dashed line in panel (a). In those years right before the bust of 2007 credit was booming as well, as shown in panel (b). The dynamics illustrated in this figure form the starting point for our empirical analysis. Panel (a) uncovers a correlation between a loose stance and future financial fragility. Figure 1 as a whole gives rise to the possibility that credit booms, which are highly predictive of crises, are triggered by too accommodative a stance of monetary policy.

In the rest of this study, we go beyond these unconditional means. We first evaluate whether the correlation between the stance and financial instability (i) remains alive and well after the inclusion of control variables, (ii) is robust to different model specifications, and (iii) reflects a causal relationship. Second, we ask if a loose stance systematically explains the dynamic illustrated in panel (b).

Outline of our empirical approach Any historical-empirical approach must recognize that financial crises are rare events. Furthermore, we face the conjecture that loose monetary policy has implications for macro-level financial stability only in the medium term since risk and leverage need time to build up. For these two reasons, we have to rely on a long span of historical data covering many economies to have sufficient sample variation to investigate our hypothesis. In addition, the natural real rate of interest is not directly observable. And so, to construct the stance of monetary policy we have to draw on state-of-the-art Bayesian methods to estimate r^* on our sample, another crucial step.

In the next section, we outline how we address these key empirical and methodological challenges. Equipped with estimates for the stance of monetary policy, section 3 then shows that loose monetary policy predicts financial crises in the medium term. We argue that overheated financial markets are a driver of this link between accommodative monetary policy and financial vulnerabilities. Evidence that our estimates uncover causal relationships is provided in section 4. Section 5 shows that the financial stability risk inherent in loose monetary policy carries over to real economic activity and section 6 concludes.

2 Framework

Our empirical analysis is based on the latest release of the Jordà-Schularick-Taylor (JST henceforth) Macrohistory Database which combines macro-financial data with a banking crisis chronology for 18 advanced economies over the period from 1870 until 2020. The database is described in Jordà, Schularick, and Taylor (2017).[3] For this study, we shall ignore the world war periods (1914–18 & 1939–45) and we also exclude the German economy during hyperinflation (1922 & 1923), but we keep all other data points of the JST Database in the analysis that follows. Our final sample has 2500 country-year observations.

Of course, not all variations in real policy interest rates are due to changes in the stance of monetary policy—some of the dynamics of interest rates are due to secular trends in the natural rate of interest arising for other fundamental reasons, such as demographics, productivity growth, or a host of other factors (see, e.g., Rachel and Summers, 2019). The natural rate is formally defined by Woodford (2003, p. 248) as the "equilibrium real rate of return in the case of fully flexible prices". If monetary policy affects real economic activity only through nominal rigidities, then "the natural rate of interest is the counterfactual rate that would be observed 'in the absence' of monetary policy. Therefore, it summarizes the real forces driving the movements in interest rates, abstracting from the influence of monetary policy decisions" (Del Negro, Giannone, Giannoni, and Tambalotti, 2017, p. 236).

[3] See `https://www.macrohistory.net/database/`. We use the 6th release of the database. See Figure A1 for an overview of included countries.

In order to isolate the stance of monetary policy from this counterfactual equilibrium rate, we need to decompose the observed ex-post real interest rate r into two latent variables, a natural rate component r^* and a residual component which we denote as *stance*. That is, we have $r \equiv r^* + stance$.[4] In this definition, "[t]he natural or 'equilibrium' real interest rate provides a benchmark for measuring the stance of monetary policy, with policy expansionary (contractionary) if the short-term real interest rate lies below (above) the natural rate" (Holston, Laubach, and Williams, 2017, p. S59). This definition of the current stance is independent of the underlying monetary policy regime, which has varied greatly across the historical eras of our study. Policy maps the state of the home (and foreign) economy into a choice of interest rate, while stance measures the implied tightness or looseness of policy, regardless of the type of control being exercised by the policymaker.

2.1 Estimating the natural rate and stance

We estimate *stance* by extending the work of Del Negro, Giannone, Giannoni, and Tambalotti (2019; henceforth DGGT). Using JST data, they estimate a long-run trend component of interest rates for 7 countries (Canada, Germany, France, Italy, Japan, the UK, and the US) by exploiting the joint dynamics of inflation and the short and long end of the yield curve. We extend their framework to all 18 countries that are covered by the JST Database.

Unlike many studies that estimate r^* (e.g., Holston, Laubach, and Williams, 2017; Fiorentini, Galesi, Pérez-Quirós, and Sentana, 2018; Jordà and Taylor, 2019), the DGGT approach explicitly accounts for a global factor in secular movements of the variables of interest.[5] DGGT identify country-specific and global trends in interest rates and inflation and corresponding stationary components in a VAR model with common trends. They do so by imposing long-run restrictions on the short- and long-ends of the yield curve. These restrictions are derived from an open-economy asset pricing model in which a marginal international investor prices bonds of all countries, implying a no-arbitrage condition in the long run. This trend-cycle decomposition is performed by Bayesian estimation.

Formally, let $R_{i,t}$ and $R^L_{i,t}$ denote the short-term and long-term nominal interest rates, respectively, and $\pi_{i,t}$ the CPI rate of inflation of country i in year t. $\bar{r}^w_t$, $\bar{\pi}^w_t$, and $\bar{ts}^w_t$ refer to world trends in the short-term real rate, inflation rate, and term spread. The latter is defined as the difference between long-term and short-term rates. $\bar{r}^i_t$, $\bar{\pi}^i_t$, and $\bar{ts}^i_t$ denote corresponding idiosyncratic trends. $\tilde{R}_{i,t}$, $\tilde{R}^L_{i,t}$, and $\tilde{\pi}_{i,t}$ are stationary components of the variables mentioned earlier. We combine trend components, stationary components, and

[4] This expression was used before by Jordà and Taylor (2019).

[5] We also replicated the Holston, Laubach, and Williams (2017) approach and extended it to all 18 countries of our study. In Appendix A2, we outline our estimation strategy, illustrate our estimated series of the natural rate, and show that the key results of this paper also hold with this alternative series of the natural rate.

observables into the column vectors $\bar{y}_t$, $\tilde{y}_t$, and y_t, respectively. With 18 countries, $\tilde{y}_t$ and y_t each contain $54 = 18 \times 3$ elements. Notice that the world trends imply that $\bar{y}_t$ has three additional rows, and thus 57 elements.

The state space representation of the model using DGGT's framework is then given by

$$\text{State equation:} \quad \begin{aligned} \bar{y}_t &= \bar{y}_{t-1} + e_t\,, \\ \tilde{y}_t &= \boldsymbol{\phi}\tilde{y}_{t-1} + \epsilon_t\,, \end{aligned} \quad \text{with} \begin{pmatrix} e_t \\ \epsilon_t \end{pmatrix} \sim N\left(\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \Sigma_e & 0 \\ 0 & \Sigma_\epsilon \end{pmatrix}\right), \tag{SE}$$

$$\text{Measurement equation:} \quad y_t = \boldsymbol{\Lambda}\bar{y}_t + \tilde{y}_t\,. \tag{ME}$$

The model is agnostic on the relationships between the three variables at all frequencies but the long term. Hence, no additional restrictions are imposed on the structure of (SE). Only the structure of $\boldsymbol{\Lambda}$ in (ME) is restricted based on the no-arbitrage conditions mentioned earlier. To be precise, it is assumed that the trends of country i's short-term real rate and term spread equal $\bar{r}^w_t + \bar{r}^i_t$ and $\bar{ts}^w_t + \bar{ts}^i_t$, respectively. The authors do not impose an equivalent to a no-arbitrage condition for inflation. Rather, they allow for country-specific loadings λ^π_i on the world inflation trend. Hence, trends in inflation are not the sum of global and country-specific trends but equal $\lambda^\pi_i \bar{\pi}^w_t + \bar{\pi}^i_t$.

The respective rows of (ME) for country i thus read as follows,

$$R_{i,t} = \bar{r}^w_t + \bar{r}^i_t + \lambda^\pi_i \bar{\pi}^w_t + \bar{\pi}^i_t + \tilde{R}_{i,t}\,, \tag{1}$$

$$R^L_{i,t} = \bar{r}^w_t + \bar{r}^i_t + \bar{ts}^w_t + \bar{ts}^i_t + \lambda^\pi_i \bar{\pi}^w_t + \bar{\pi}^i_t + \tilde{R}^L_{i,t}\,, \tag{2}$$

$$\pi_{i,t} = \lambda^\pi_i \bar{\pi}^w_t + \bar{\pi}^i_t + \tilde{\pi}_{i,t}\,. \tag{3}$$

Then, following DGGT, we will interpret $r^*_{i,t} \equiv \bar{r}^w_t + \bar{r}^i_t$ as the natural rate of interest of country i.

After subtracting equation (3) from (1), we have

$$\underbrace{R_{i,t} - \pi_{i,t}}_{r} = \underbrace{r^*_{i,t}}_{r^*} + \underbrace{\tilde{R}_{i,t} - \tilde{\pi}_{i,t}}_{stance}\,,$$

which is the aforementioned decomposition of the observed ex-post real interest rate into the natural rate trend component and the residual stance component that can be controlled by a country's central bank.

Here, the terms $\tilde{R}$ and $\tilde{\pi}$ will capture all stationary short-term and medium-term deviations from long-run trends in short-term nominal rates and inflation, respectively. We should remark that, since the DGGT trend-cycle decomposition abstracts from relationships between variables at all frequencies but the long run, $\tilde{R}$ and $\tilde{\pi}$ have no economic interpretation. We therefore will not analyze separately these two terms in further detail in the rest of this paper, and instead focus only on *stance* itself.

We are interested in the impact of persistently loose monetary policy rather than in single periods of undershooting.[6] Therefore, we base our analysis on a moving-average transformation of *stance* to better capture sustained periods of loose policy. We construct a lagged-average stance of monetary policy of country i in year t defined as the five-year averaged deviations of the ex-post real interest rate from its natural rate counterpart,

$$\overline{stance}_{i,t} = \frac{1}{5}\sum_{k=0}^{4}\left(r_{i,t-k} - r^{*}_{i,t-k}\right) . \tag{4}$$

This stance measure is our main object of interest throughout this study. It represents the relevant independent variable in all econometric models we estimate. Robustness checks reported later show that the exact number of years over which we take moving averages does not affect our results significantly.

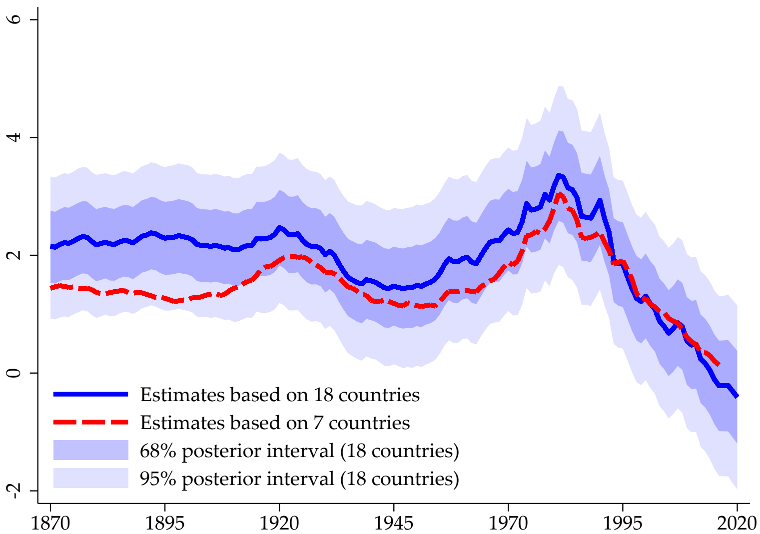
Regarding the estimation procedure, we try to follow as closely as possible the approach in DGGT. Our sample includes a larger number of countries which constitutes a more heterogeneous group compared with the 7-country model of DGGT. It is thus no surprise that our estimated series of $\bar{r}^{w}_{t}$ would exhibit a larger degree of variation than the replicated series of DGGT if we used the same priors as DGGT. We therefore tighten the priors slightly.[7] Some further work is needed to handle missing data. As done by DGGT and explained in footnote 21 of their paper, for the estimation of this model we ignore observations for which any of the three observed variables exceeds 30% in absolute values. Data for short-term rates, long-term rates, and inflation is then missing at the same time only for pre-1920 Ireland, the war-torn economy of Belgium (1915–1919), and Germany during hyperinflation (1923). We ignore pre-1920 Ireland and, following DGGT, we interpolate over all other missing values. We take 100,000 draws and treat the first 50,000 as the burn-in sample.

Figure 2 shows estimated world trends of the natural rate of interest, that is posterior medians of $\bar{r}^{w}_{t}$, for both our model with 18 countries and the original DGGT model with

[6] When departing from the standard rational expectations assumption and introducing sticky expectations about inflation, the real interest rate can deviate from the natural rate for many years. Bianchi, Lettau, and Ludvigson (2022) introduce a model that features such long-lasting deviations from monetary neutrality after regime changes in the conduct of monetary policy.

[7] We maintain the assumption that the prior for the covariance matrix Σ_e of equation (SE) is an inverse Wishart distribution with κ_e degrees of freedom. However, we raise κ_e from 100 to 200. Furthermore, we set the prior for Σ_e to have a mode equal to a diagonal matrix with elements equal to 0.007 instead of 0.010 for interest rates, and to 0.014 instead of 0.020 for inflation. Similar adjustments were made by Cesa-Bianchi, Harrison, and Sajedi (2022) who replicate the DGGT approach for a sample of 31 countries.

Figure 2: *Trends in the global real rate.*



Notes: The figure shows posterior medians of $\bar{r}_t^w$ with corresponding posterior coverage intervals for our model with 18 countries (which uses data until 2020) and posterior medians of $\bar{r}_t^w$ for the original DGGT model with 7 countries (which uses data until 2016).

7 countries. The two estimated series closely resemble each other in the later part of the sample period. For the pre-1970 period, we estimate a higher posterior median of $\bar{r}_t^w$. As noted by DGGT, since the late 1970s "the U.S. trend *is* the global trend" implying smaller country-specific trend components. This finding carries over to our 18-country model; before the late 1970s, country-specific trends were larger.[8] The inclusion of the 11 additional countries shifts the world trend upward in the period before the late 1970s. The reason is that these 11 countries had higher real interest rates on average.[9] We can note in this context that Cesa-Bianchi, Harrison, and Sajedi (2022), who estimate the DGGT model for 31 countries over the period from 1900 to 2015, also obtain higher posterior medians for the world trend in the first part of their sample compared to the original DGGT estimates.

We also highlight the unprecedentedly low natural rates at the end of the sample. This is a finding that also emerges in other studies that estimate the natural rate of interest (Holston, Laubach, and Williams, 2017; Jordà and Taylor, 2019). According to our definition of stance in equation (4), many countries were therefore not in an environment characterized by exceptionally loose monetary policy in the 2010s.

Full country-by-country estimates are shown in the Appendix. Figure A1 shows esti-

[8]In the post-1977 period, the cross-country average of the absolute value of the idiosyncratic trend component $\bar{r}_t^i$ is 0.40. The corresponding average for the period up to 1977 is 0.47. Also see Figure A1.

[9]In the post-1977 period, the cross-country average of real interest rates is 1.63% for the G7 and 1.65% for the other 11 countries. Corresponding averages for the period up to 1977 are 1.88% and 2.50%.

mated country trends of the natural rate of interest, that is posterior medians of $\bar{r}_t^w + \bar{r}_t^i$, as well as $\overline{stance}_{i,t}$ as defined in equation (4) for all 18 countries of our sample. Whenever applicable, we also show replicated country trends from the 7-country DGGT model.

To sum up, extracting the trend matters. Today's investors live in a low real rate world, but in the past their ancestors did not. For example, the red line in the bottom-right panel of Figure A1 shows that an American investor faced a high interest rate environment in the 1870s. However, our estimates show that it was not as high as suggested by the ex-post real interest rate. While the latter was over 8% (in the JST data), the average value of $r_{USA} - r^*_{USA}$ was only 5.88% in the 1870s due to high natural rates (blue line). In contrast, while $r_{USA} - r^*_{USA}$ was on average negative in the 2010s (–0.98%), those values are by no means unprecedented. In other words, a Millennial investor might not behave excessively riskily in this era of low real rates, once we make allowance for the very different reference level of the natural rate they had to live with.

2.2 Statistical design

In the remainder of this study, we analyze different questions within the same econometric model. In its general form, this linear probability model with a binary outcome B can be written as a Jordà (2005) local projection, or LP,

$$B_{i,t+h} = \beta^h \, \overline{stance}_{i,t} + \alpha_i^h + \mathbf{\Gamma}^h \, \mathbf{X}_{i,t} + \sum_{k=1}^{5} \delta_k^h \, b_{i,t-k} + u_{i,t+h} \,, \qquad h = 0, \ldots, H \,. \tag{5}$$

We will first estimate the model by OLS before validating our results with an IV approach. The linear probability model has the advantage that we can interpret directly the coefficients in terms of changes in the outcome probability (a logit specification is also explored for robustness; it is more cumbersome but yields very similar results).

In this LP-OLS setup, the outcome variable $B_{i,t}$ is a binary variable that captures, depending on the context, events of financial turmoil (sections 3.1 and 4.2), periods of overheated credit and asset markets (sections 3.3 and 4.3), or years of sharp downturns in real economic activity (section 5). It is equal to 1 when country i experiences such an event and 0 otherwise. The lagged terms $b_{i,t-k}$ capture past information about the dependent variable, as defined below. While B, and thus also b, takes on different forms in the following sections, everything else in model (5), including the set of control variables, remains unchanged throughout the rest of the paper.

The intercept terms α_i^h are country fixed effects at horizon h. $\mathbf{X}$ is a vector of standard macro-financial variables that controls for confounding factors that are potentially correlated with both $\overline{stance}$ and the binary outcome variable such as the economy's position in the business cycle. It includes annual changes from year $t-5$ to t of log-transformed real GDP per capita, log-transformed consumer prices, the log-transformed local currency's price vis-à-vis the US-Dollar, and the investment-to-GDP and credit-to-GDP ratios. It also contains the natural rate of interest and the slope of the yield curve as of period t, that is r_t^* and $R_t^L - R_t$. Including r_t^* is an additional guarantee that we capture the economy's position in the business cycle. Furthermore, this variable also controls for the general interest rate environment of the economy; since the stance of monetary policy might affect the natural rate (Hillenbrand, 2021; McKay and Wieland, 2021; Kashyap and Stein, 2023) or vice versa and could be a potential source of financial instability on its own, it is a possible confounding factor. Similarly, monetary policy affects the slope of the yield curve, which in turn is a strong short-term predictor of financial crises (Bluwstein, Buckmann, Joseph, Kapadia, and Şimşek, 2023; Parker and Schularick, 2021).

Finally, note that we also control for the global debt-to-GDP ratio and global capital and non-core funding ratios of banks in year t. We define these global variables as unweighted averages across countries. They are a parametrically economical way to control for cross-country factors, such as global business cycle dynamics and the structure of the financial system. We also provide estimates based on controlling for global factors with time fixed effects. The results are broadly the same, though the time fixed effects results involve estimating a much larger number of parameters, thus affecting the precision of the estimates.

In the end, the β^h will be our key coefficients of interest. They trace out the predictive power of the stance of monetary policy for financial instability and credit and asset market overheating. To provide a clearer indication of the economic relevance of our estimates, we present them relative to the unconditional probability that the outcome variable is equal to 1. To ensure validity of our statistical inference, we allow for complex structures of the error term by computing Driscoll-Kraay (1998) standard errors that are robust to heteroskedasticity as well as to spatial and temporal correlation up to $ceiling(1.5 \times h)$ annual lags.

2.3 Estimation uncertainty in r^*

Since the natural rate r^* is a latent variable, our estimation procedure consists of two stages. First, we estimate r^* as described above to construct our measure of the stance of monetary policy, $\overline{stance}$. In the second stage, we use $\overline{stance}$ as a regressor in the local projections. Our estimates of r^*—and consequently of $\overline{stance}$—are subject to uncertainty, as illustrated in Figure 2. This additional source of estimation uncertainty could create two problems, which we address as follows.

One issue that can arise from the two-stage estimation procedure is that the first-stage uncertainty may not be reflected in the analytic standard errors, leading to inaccurate statistical inference. To address this concern, we calculate alternative confidence bands for our main results using the panel moving blocks bootstrap method proposed by Gonçalves (2011), which resamples contiguous rows of data without affecting the cross-sectional structure of the data. In order to take the first-stage uncertainty into account, we extend this method along one dimension. Specifically, we create rows by combining the data and a random draw from the 50,000 posterior draws of r^*. Each bootstrap sample is created by drawing, with replacement, a different posterior draw of r^*. Apart from this extension, we apply the bootstrap method in the usual way. We choose a block length of five years and compute bootstrap confidence intervals with asymptotic refinement (Cameron and Trivedi, 2005, p. 364) based on 1,000 bootstrap samples.[10]

A second potential issue is that the first-stage uncertainty is akin to a measurement error that biases the second-stage coefficient estimates towards 0. The LP-OLS estimates presented in the next section demonstrate a significant association between a loose stance and the risk of future financial instability. However, they might be biased towards 0 and should, therefore, be interpreted as lower bounds. In section 4, we introduce an LP-IV setup that addresses this measurement error problem. Indeed, the IV coefficient estimates are larger, as expected, suggesting an even stronger relationship between an accommodative stance of monetary policy and the likelihood of financial turmoil down the road.

3 Loose monetary policy and financial instability

3.1 Financial crises

Are periods of persistently loose monetary policy more crisis-prone? This section argues that the answer to this question is in the affirmative.

Turning to define our main outcome of interest, let $crisis_{i,t}$ be a dummy that is equal to 1 if a financial crisis starts in country i in year t. It is zero otherwise. In this part, we estimate model (5) with: $B_{i,t} = \max\{crisis_{i,t}, crisis_{i,t+1}, crisis_{i,t+2}\}$;[11] and $b_{i,t} = crisis_{i,t}$. That is, for the definition of B, we consider crisis risk over a three-year window. This definition of the dependent variable B captures the idea that it is hard to identify the exact starting year of financial crises but easier to pinpoint "danger zones" in which crisis risk is elevated (Schularick, ter Steege, and Ward, 2021). We will ensure that results hold for alternative

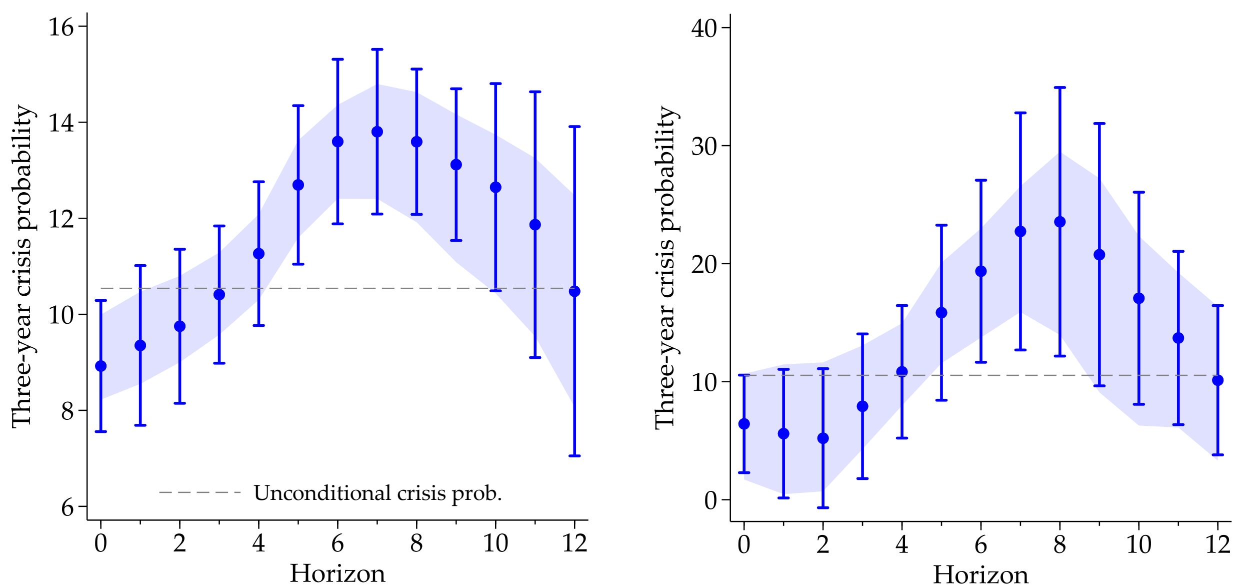
[10] We show bootstrap confidence intervals for our main results on the relation between the *stance* and crisis risk (Figures 3 and 11). We also verified the robustness of our other results to this first-stage estimation uncertainty.

[11] There is never more than one financial crisis within a 3-year horizon in our sample.

Figure 3: *The connection between loose monetary policy and financial crises.*

(a) $\overline{stance}$ *lower by 100 bps*

(b) $\mathbb{1}\{\overline{stance} < 20^{th}\ percentile\} = 1$



Notes: Panel (a) shows estimates of $\{-100\beta^h\}_{h=0}^{12}$ of equation (5) with B and b as defined in the text relative to the unconditional full-sample three-year crisis probability. Panel (b) replaces the continuous variable $\overline{stance}_{i,t}$ by the binary variable $\mathbb{1}\{\overline{stance}_{i,t} < 20^{th}\ percentile\}$ and shows estimates of $\{100\beta^h\}_{h=0}^{12}$ relative to the unconditional three-year crisis probability. Control variables are outlined in section 2.2. Bars indicate 95% confidence intervals based on Driscoll-Kraay (1998) standard errors with $ceiling(1.5 \times h)$ lags. The shaded areas denote 95% confidence intervals based on the bootstrap procedure described in part 2.3.

horizons of this danger zone variable. In particular, we show below that loose monetary policy even predicts the *exact* starting year of financial crises with a high degree of statistical precision. That is, our main result also holds for $B_{i,t} = b_{i,t} = crisis_{i,t}$.

Figure 3 presents our main result. Panel (a) shows the measured change in crisis risk, relative to the full-sample unconditional three-year crisis probability (dashed line), when $\overline{stance}$ as defined in equation (4) is lower by one percentage point. Panel (b) displays the rise in the likelihood of crises when $\overline{stance}$ is in the lowest quintile. Bars indicate corresponding 95% confidence intervals based on analytic standard errors. The shaded areas represent bootstrap confidence intervals and ensure that the first-stage uncertainty discussed in part 2.3 does not have a significant impact on statistical inference.

Let us focus on panel (a) first. We see significant estimates in the medium term, that is around horizons of 5 to 10 years.[12] Financial crises are predicted by loose monetary

[12]The "short term" and "medium term" are of course not clearly defined notions. One might wonder why we interpret horizons up to 4 years as the "short term". Note in this context that the amplitude of the financial cycle is much longer than for the regular business cycle (Borio, 2014).

policy several years ahead. The importance of this empirical finding does not only arise from its high level of statistical significance and—as we will see below—robustness to model specification, but also from economic relevance. If the averaged monetary-policy-determined part of real rates is 1 pp lower, then financial crisis risk increases by 2.2 pps 5 to 7 years ahead and by 3.3 pps 7 to 9 years ahead. Since the unconditional probability that a financial crisis starts within a 3-year window is only 10.5% in our full sample as indicated by the dashed line, these estimates are of a nontrivial magnitude.

Notice that the sample mean of $\overline{stance}$ is approximately 0. Hence, when monetary policy is neutral ($\overline{stance} = 0$) and all other covariates are at their sample means as well, three-year crisis risk is roughly 10.5%. Our estimates suggest that this probability rises to about 17.1% when $\overline{stance} = -2.5\%$ (which is the 20^{th} percentile of the pooled country-year $\overline{stance}$ distribution) 6 to 8 years ahead.

We do not find evidence for a positive link between loose monetary policy and financial vulnerabilities in the short term. If anything, point estimates in Figure 3 indicate a *negative* relation between financial fragility and a loose stance at horizons below 4 years. This is in line with the literature. We already mentioned in the Introduction Wicksell's (1898) line of reasoning that low interest rates spur house price growth and open the door for speculation and boom-and-bust cycles. He also hypothesized in this context that *during* the boom, the price of credit will actually rise (p. 90). In the theoretical models of Boissay, Collard, Galí, and Manea (2022) and Akinci, Benigno, and Del Negro (2020), financial crises are preceded by a long period of low interest rates *and an abrupt reversal of rates*. On the empirical side, Demirgüç-Kunt and Detragiache (1998) document that high real interest rates are associated with a contemporaneous higher probability of banking crises. Schularick, ter Steege, and Ward (2021) and Jiménez, Kuvshinov, Peydró, and Richter (2022) find that "last minute" leaning against credit booms is more likely to trigger rather than avert financial crises. A point in case is the Fed's "restrictive action early in 1928" with the intention "to curb the stock market boom" (Friedman and Schwartz, 1963, p. 289). Furthermore, as already noted above, there is strong evidence that a flattening yield curve predicts financial crises. Before such crises, the short end of the yield curve rises while the long end stays put (Parker and Schularick, 2021). One explanation for this finding is the emergence of a credit and asset price boom in the immediate years before financial crises. The empirical and theoretical literature outlined in the Introduction suggests that loose monetary policy could trigger such overheated credit and asset markets. We provide evidence in favor of this link below.

Is a simple indicator that captures periods of *very loose* monetary policy able to predict financial instability? We address this question by replacing the continuous stance variable by a binary one in model (5). This binary variable is equal to 1 if $\overline{stance}_{i,t}$ is in the lowest quintile of its full-sample pooled country-year distribution and 0 otherwise. Estimated coefficients of interest of this modified model are shown in panel (b) of Figure 3. Impulse responses exhibit a similar pattern as in panel (a); too accommodative a stance again predicts a higher likelihood of banking crises in the medium term, starting at a horizon of 6 years. Point estimates are large. At the peak, a central bank that is very accommodative from a historical point of view (using this indicator) increases the probability of entering a financial crisis danger zone by 13 pps.

Our main result that loose monetary policy predicts financial instability is based on the narrative banking crisis chronology of the JST Database (hereinafter JST chronology). However, it is not dependent on this historical account of financial crises. A similar pattern is observable for alternative notions of financial vulnerabilities. This is shown in Figure 4 using the continuous measure of stance. Panels (a) and (b) in this figure estimate the same model for two alternative banking crisis chronologies constructed by Reinhart and Rogoff (2009) and Baron, Verner, and Xiong (2021) and display a similar picture.[13] In panel (c), we replace the crisis indicator with the bank equity crash indicator from Baron, Verner, and Xiong (2021). This binary indicator flashes red in years of real equity return declines above 30%. It covers all 18 countries of the JST Database over the period from 1870 until 2016. The estimates suggest that the likelihood of bank equity crashes increases by more than 2 pps 4 to 6 years ahead when $\overline{stance}$ is 1 pp lower. This measure remains elevated in subsequent years and peaks at 5.3 pps at a horizon of $h = 8$ years. Again, these are not only statistically significant but also economically meaningful estimates given that the unconditional three-year bank equity crash probability in our sample is only 12%, as indicated by the dashed line.

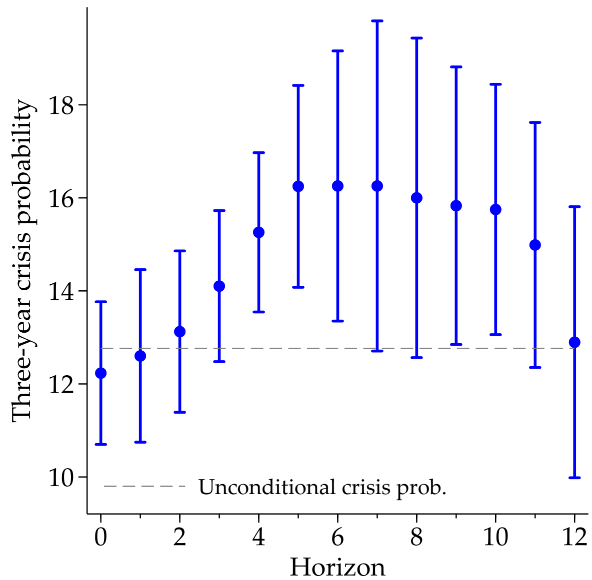
Recall that our dependent variable is defined as $B_{i,t} = \max\{crisis_{i,t}, crisis_{i,t+1}, crisis_{i,t+2}\}$. That is, we consider crisis risk over three years as explained in the beginning of this section. We can re-define the dependent variable as $B_{i,t} = \max\{crisis_{i,t}, crisis_{i,t+1}, \ldots, crisis_{i,t+F}\}$ and re-estimate model (5) for different values of F. Results of this exercise are shown in Figure A2. We obtain significant estimates even for $F = 0$, starting at a horizon of 6 years. That is, we can also predict the *exact* starting year of financial crises at the 95% level. Similarly, our main result shown in Figure 3 also holds when averaging $r - r^*$ over a different number of years. This is illustrated in Figure A3.

Throughout this study, our findings are derived from the estimation of a simple linear probability model. This simplifies interpretability but has obvious shortcomings such as possible out-of-bounds predicted probabilities. Allowing for a non-linear relationship

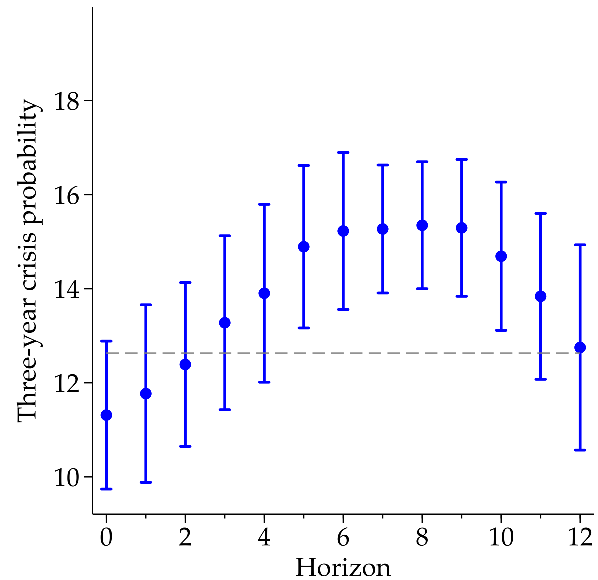
[13] For a comparison between the JST chronology and the BVX alternative chronology, see the Documentation of the former which can be found here: `https://www.macrohistory.net/database/`. See also the discussion of chronologies in Sufi and Taylor (2022).

Figure 4: *The connection between loose monetary policy and alternative financial instability event indicators.*

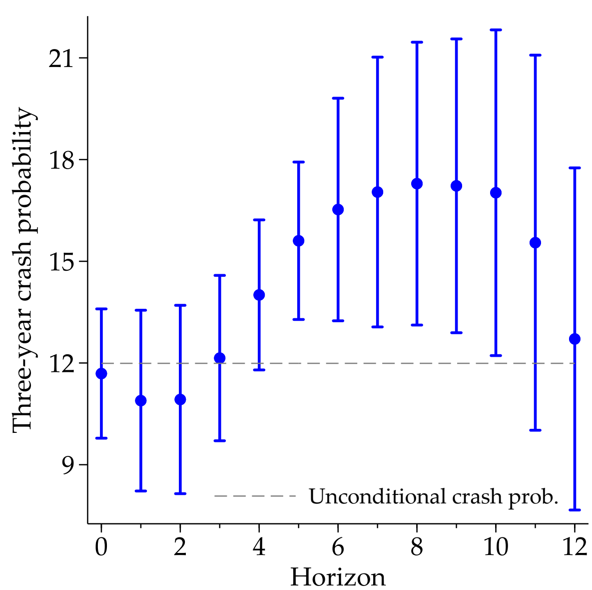
(a) *Baron et al. (2021) financial crises,*
$\overline{\text{stance}}$ lower by 100 bps



(b) *Reinhart and Rogoff (2009) financial crises,*
$\overline{\text{stance}}$ lower by 100 bps



(c) *Baron et al. (2021) bank equity crashes,*
$\overline{\text{stance}}$ lower by 100 bps



Notes: The same notes as in Figure 3 apply.

between crisis probability and regressors does not have much impact on our results. Figure A4 reveals a similar pattern within a logistic model.

One possible concern is that a low stance simply reflects periods of economic expansions. If recessions—some of which are due to financial crises—follow expansions after a roughly fixed time span, then our main results could just describe a textbook-like real business cycle. This critique is unjustified for two reasons. First, we control for the position in the (local and global) business cycle by including a rich set of macro-financial control variables as outlined above. Our results show that a loose stance predicts financial crises *irrespective of* the state of the business cycle. Second, we can show that $\overline{stance}$ is not a signal for normal business-cycle downturns; it only predicts recessions associated with financial turmoil.

Evidence for this latter claim is provided in Figure A5. There, we re-estimate our main econometric model (5) for a different dependent variable. This dependent variable is a binary indicator for one of two mutually exclusive and exhaustive types of recessions defined as in Jordà, Schularick, and Taylor (2016); those that are associated with financial crises and those that are not. A low value of $\overline{stance}$ does not measure a higher medium-term probability of normal recessions. It predicts only those recessions that go hand-in-hand with a systemic banking crisis.

We take global factors into account by including global variables in our set of controls. We obtain similar results when we enrich our set of control variables with decade fixed effects or when we replace the global variables with decade fixed effects. This is shown in panels (a) and (b) of Figure A6, respectively.

Finally, the relationship between the stance of monetary policy and financial instability risk remains alive and well for the post-WWII period (Figure A7) and when using the Holston, Laubach, and Williams (2017) approach to estimate the natural rate of interest (Figure B2). Alas, an extension of the HLW approach to all countries and years of our dataset is non-trivial. Sensitive adjustments are necessary as outlined in Appendix A2, which in turn are reflected in the heightened estimation uncertainty apparent in Figure B2. Nevertheless, point estimates exhibit a similar shape as in Figures 3 (a) and 4.

3.2 The relevance of the stance of monetary policy

Our stance measure seeks to capture movements in only one component of the interest rate that is affected by monetary policymakers. One might wonder why we focus on this component of interest rates rather than on the absolute level of the interest rate, as a whole. So far, we showed that the stance predicts financial instability. One question that remains, though, is the role of the level ex-post real interest rate. In this part, we argue that this level is not predictive of crises. Rather, the residual stance is what matters, i.e., "discretionary loose monetary policy may on its own be a source of financial instability" (Boissay, Collard, Galí, and Manea, 2022).

As outlined in the Introduction, from a behavioral point of view we can interpret the natural interest rate as a history-dependent reference rate. Hence, the stance captures agents' perceived level of the current interest rate environment relative to that reference.

Leaving this behavioral view behind, recall that r^* equates savings and investments in an economy without nominal rigidities. If interest rates are lower than r^*, then the economy generates excess liquidity. In other words, when monetary policy pushes rates below r^*, real investment opportunities exceed real savings. Such an accommodative stance of monetary policy might pave the way for debt- and leveraged-financed investment booms. Investment booms and a consequential capital overhang due to too loose a stance of monetary policy are precursors of financial crises in the model of Boissay, Collard, Galí, and Manea (2022).

The natural rate r^* is the long-run trend of interest rates. We can therefore also think of r^* as a proxy for rates on long-term fixed-interest liabilities of financial institutions. When observed rates are lower than r^*, financial institutions with a large share of such long-term liabilities are obliged to search for yield (Rajan, 2005).

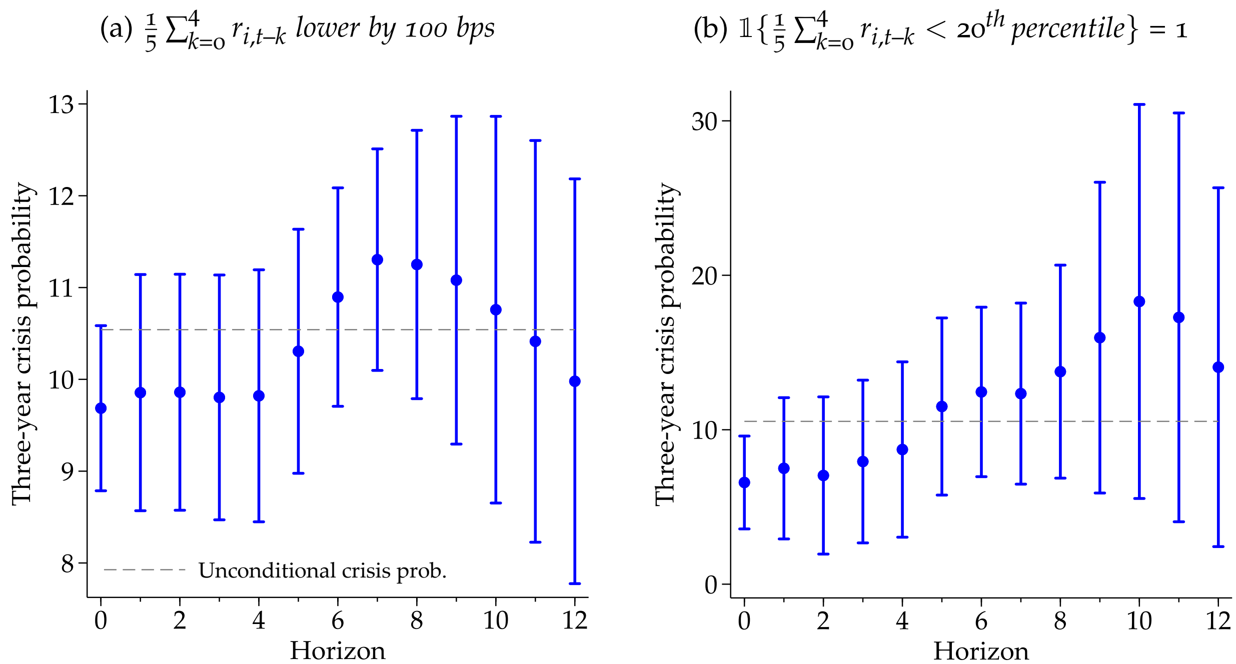
Also note that real interest rates have been on a steady downward trend over many centuries (Schmelzing, 2020; Jordà, Singh, and Taylor, 2022). Yet, financial crises have always been around. From a long-term historical perspective, it is therefore hard to argue for the relevance of the absolute level of real interest rates as a source of financial instability.

From a methodological point of view, when analyzing the consequences of low interest rate environments, the literature seldom distinguishes between secular trends in interest rates and shifts in the stance of monetary policy (Boyarchenko, Favara, and Schularick, 2022). Our framework allows us to do exactly that.

Figure 5 further illustrates the predominant importance of the stance of monetary policy for financial stability considerations, rather than the absolute level of rates. It shows estimates of model (5) when the only interest rate variable included in the model is the level measure $\frac{1}{5}\sum_{k=0}^{4} r_{i,t-k}$. The outcome variable refers to JST financial crises, defined as before. The figure indicates that real interest rates *on their own* are absolutely *not* informative for crisis risk down the road. Neither the full distribution of real interest rates in panel (a), nor its left tail in panel (b), is associated with higher medium-term financial stability risk.

Taken together, the strong, significant, and long-lasting link between loose monetary policy and financial instability presented in Figure 3 vanishes when considering only the absolute level of real interest rates separately. This strongly suggests that it is the conduct of monetary policy rather than interest rate levels as such that plays the dominant role in our main results.

Figure 5: *The connection between real interest rates and financial crises.*



Notes: The same notes as in Figure 3 apply with one difference; the only interest variable included in model (5) is now $\frac{1}{5}\sum_{k=0}^{4} r_{i,t-k}$.

3.3 Credit and asset market overheating

Framework So far, we have shown that the monetary policy stance predicts financial instabilities in the medium term, suggesting that excessively loose monetary policy sets off dynamics in financial markets that often end up badly. In this section, we get to the bottom of the question of what these dynamics are.

In the Introduction, we outlined financial variables—rising leverage and asset prices—that have been identified by the existing empirical macro-finance literature as important harbingers of financial crises (Jordà, Schularick, and Taylor, 2015b). It is precisely the dynamics in these variables that the theoretical literature connects to loose monetary policy. We now go one step back and ask if loose monetary policy triggers such unsustainable trends in financial markets.

More precisely, our aim here is to build on the recent findings of Greenwood, Hanson, Shleifer, and Sørensen (2022). They define Red-zones or *R-zones* in which both credit growth and asset price growth are elevated. The authors show that these R-zones have a crucial impact on the stability of the financial system. R-zone signals have a high degree of predictability for financial crises that goes far beyond the predictive power inherent in credit growth alone.

Table 1: *Relevant percentiles of private debt and asset price changes.*

	Post–1949 sub-sample	Full sample
80^{th} *percentile of* $\Delta_3 100\,(Debt/GDP)$		
Household credit	6.23	6.12
Business credit	4.73	4.69
66.7^{th} *percentile of* $\Delta_3 100\,(\log Price)$		
House prices	12.84	11.33
Stock prices	22.82	22.22

Is a loose stance of monetary policy associated with a higher probability of ending up in such R-zones? We answer this question by making use of the same framework as before. We start again from model (5), keeping the vector of control variables $\mathbf{X}$ unchanged. In the previous section, the binary dependent variable B referred to financial crisis risk. Now, it refers to R-zones: $B_{i,t} = R\text{-}zone_{i,t}$. R-zones are defined as in Greenwood, Hanson, Shleifer, and Sørensen (2022), with

$$\begin{aligned} &High\text{-}Debt\text{-}Growth_{i,t} = \mathbb{1}\{\Delta_3\,(Debt/GDP)_{i,t} > 80^{th}\ percentile\}\,, \\ &High\text{-}Price\text{-}Growth_{i,t} = \mathbb{1}\{\Delta_3\,(\log Price_{i,t}) > 66.7^{th}\ percentile\}\,, \\ &R\text{-}zone_{i,t} = High\text{-}Debt\text{-}Growth_{i,t} \times High\text{-}Price\text{-}Growth_{i,t}\,. \end{aligned} \tag{6}$$

Here, the High-Debt-Growth indicator takes on a positive value if the three-year change in the debt-to-GDP ratio is in the top quintile of the pooled country-year debt-growth distribution. The High-Price-Growth indicator is defined in a similar way for the top tercile. By definition, the economy is then in an R-zone when both debt and prices rise relatively strongly from a historical point of view.

As in Greenwood, Hanson, Shleifer, and Sørensen (2022), we construct a household sector R-zone based on household credit growth and real home price growth and a business sector R-zone based on business credit growth and real equity price growth. That is, *Debt* refers to either total loans to households or to total loans to businesses, while *Price* denotes real (CPI-deflated) house prices or real stock prices, respectively. Greenwood, Hanson, Shleifer, and Sørensen (2022) find that a country that is in the household (business) sector R-zone faces a three-year crisis risk of 37 (45) percent.

Table 1 provides an overview of the relevant percentiles of the 3-year changes in the private debt and asset price variables for the household and business sectors. For example, in the post-WWII period, the High-Household-Debt-Growth indicator equals 1 if the three-year change in the household-debt-to-GDP ratio surpasses 6.23%. The corresponding threshold for business credit is lower (4.73%), in line with the finding that the increase in credit-to-GDP ratios across countries after WWII was primarily driven by the household sector (Jordà, Schularick, and Taylor, 2017; Müller and Verner, 2023).

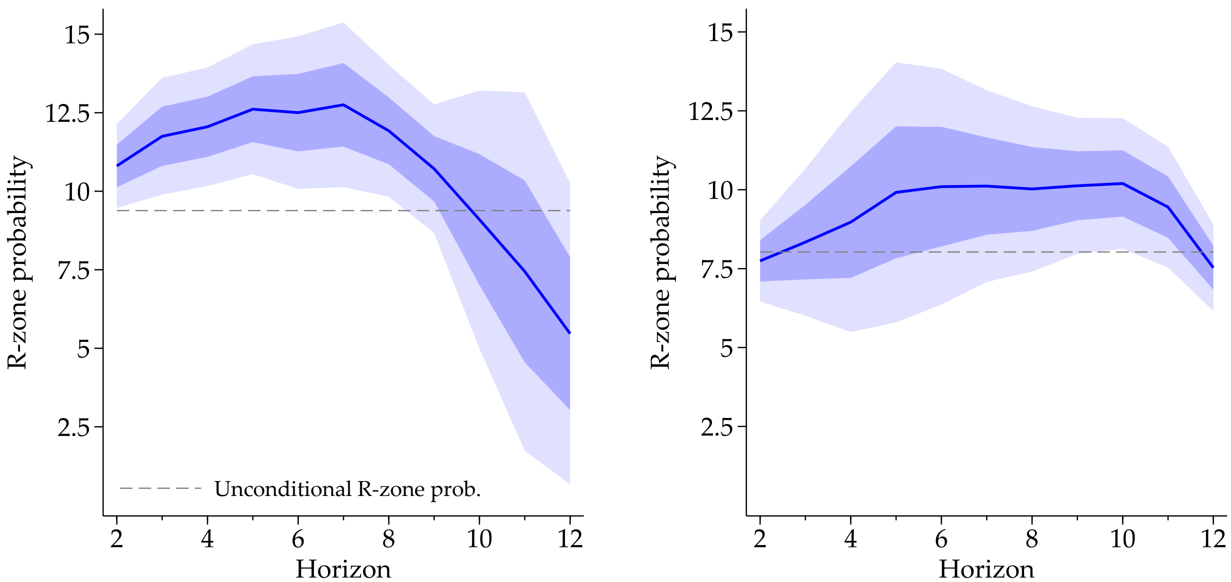
We once again control for past information about the dependent variable. However, this time, the binary outcome variables are derived from continuous macro-financial variables. Consequently, we can directly include annual changes in these variables from lag 1 to 5 as control variables. For instance, the household sector R-zone indicator is based on the growth of household loans and house prices. Therefore, when considering household sector R-zones, we control for annual changes in the household credit-to-GDP ratio and in log-transformed real house prices from lag 1 to lag 5. Apart from the re-definition of the binary outcome variable, and the adjustment in the lag outcome controls, we use the same framework as before. In particular, we use the same set of other control variables and fixed effects. Notice that the occurrence of an R-zone in year t is the outcome of an acceleration in financial prices and quantities over the preceding three years. If the R-zone indicator turns on, credit and asset price growth is already fully underway. We thus estimate the model starting at a horizon of $h = 2$ years. The rest of this section provides evidence for a relationship between the stance of monetary policy and credit and asset market overheating across time and space. In this, we follow Greenwood, Hanson, Shleifer, and Sørensen (2022) along two other dimensions. First, we focus on the same period, namely the post-WWII era. Second, we estimate the model by OLS. Further below, we extend the sample period to all available data points of the JST Database and provide IV estimates.

Results Recall that $\overline{stance}$ predicts financial turmoil at statistically significant levels starting at a horizon of around five years, as we saw in Figure 3. Panels (a) and (b) of Figure 6 now show that, at this point, it is likely that the country has already experienced financial fragility by entering an R-zone. A loose stance of monetary policy predicts the emergence of credit and asset market overheating in post-WWII advanced economies both in the household and in the business sector. That is, the stance has predictive power for those indicators that have been identified by Greenwood, Hanson, Shleifer, and Sørensen (2022) as important short-term early warning signals for crisis risk.

When stance is looser by 1 pp, the likelihood of entering a household or business sector R-zone in five years is higher by 3.2 and 1.9 pps, respectively. That is, we predict exaggerated growth in key financial quantities and prices three to five years ahead that eventually activate warning signals. Recall that we use a linear model and the sample average of $\overline{stance}$ is approximately 0. This makes point estimates easy to interpret. The unconditional mean probability of household sector R-zones, for instance, is 9.4% in the post-1949 sub-sample as shown by the dashed line in panel (a) of Figure 6. Thus, panel

Figure 6: *The connection between loose monetary policy and post-WWII R-zones.*

(a) *Household sector R-zone,* $\overline{\textit{stance}}$ *lower by 100 bps* (b) *Business sector R-zone,* $\overline{\textit{stance}}$ *lower by 100 bps*

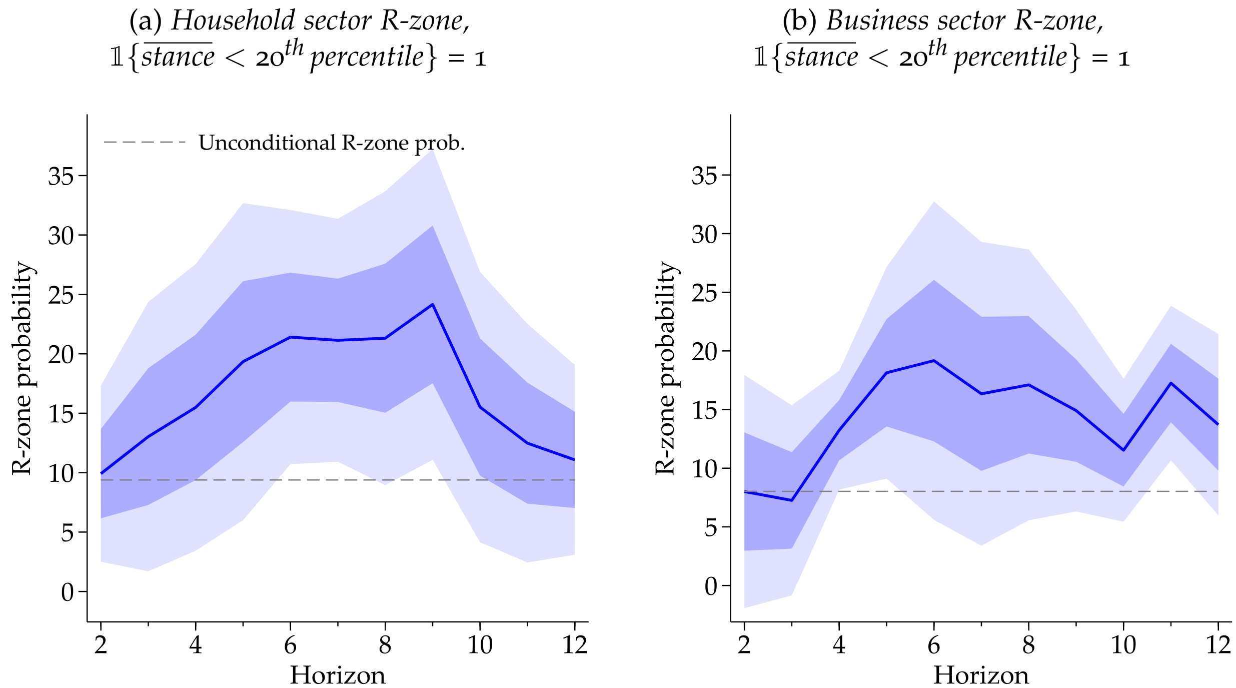


Notes: The figure shows estimates of $\{-100\beta^h\}_{h=2}^{12}$ of model (5) with R-zones as defined in equation (6) as the dependent variable relative to the unconditional post-1949 R-zone probability. The control variables are the same as before and outlined in section 2.2, and b again controls for past information about the dependent variable as described in the text. Only data for the post-1949 period is used. Shaded areas indicate 95% (light) and 68% (dark) confidence intervals based on Driscoll-Kraay (1998) standard errors with *ceiling*$(1.5 \times h)$ lags.

(a) indicates that when all covariates are at their sample means but $\overline{\textit{stance}}$ is at its lowest quintile (–2.4%), the household sector R-zone indicator turns on with a probability of 17.1% in five years, almost twice the sub-sample mean. While the association between loose monetary policy and household sector R-zones is significant at the 5% level, the level of statistical uncertainty rises when examining R-zones in the business sector. This result can be attributed to the heterogeneity within the corporate sector. Credit booms are characterized by a reallocation of credit from the tradable to the non-tradable sector, and only credit expansions to the non-tradable sector are associated with a heightened risk of financial instability (Müller and Verner, 2023). The findings presented in Figure 6 also continue to hold when we enrich the model with decade fixed effects or when we replace global control variables with time fixed effects (Figure A8). Furthermore, the estimation of a logit model yields similar results (left panels of Figure A9).

Similar to our earlier exercise with crisis prediction (panel (b) of Figure 3), we can also establish a relation between a simple indicator for *very loose* monetary policy and R-zone

Figure 7: *The connection between very loose monetary policy and post-WWII R-zones.*



Notes: The same notes as in Figure 6 apply with one difference; the continuous variable $\overline{stance}$ is replaced by the binary variable $\mathbb{1}\{\overline{stance} < 20^{th}\ percentile\}$ and estimates of $\{100\beta^h\}_{h=2}^{12}$ are shown.

risk. This is shown in Figure 7. Here again, point estimates are economically large. For instance, the figure suggests that five years after monetary policy was very accommodative from a historical point of view, the probability of entering a business sector R-zone increases from 8% to 18.1% and remains high in the following years. This suggests strong growth in business credit and equity prices at a horizon from 3 to 5 years.

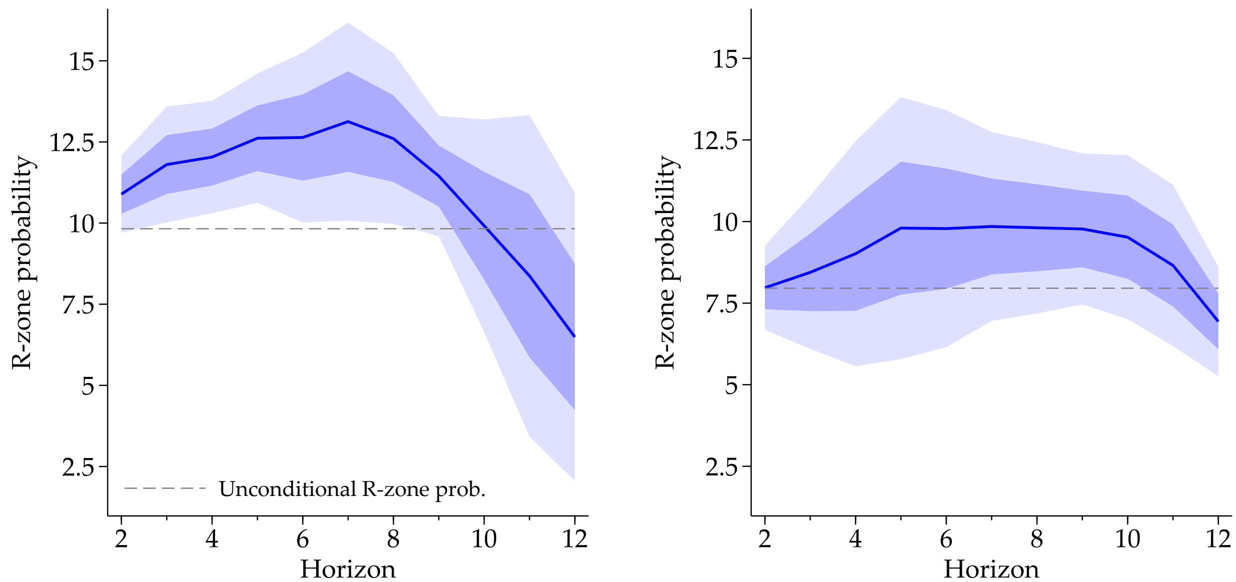
So far, we restricted our sample to the post-1949 era, the same period over which Greenwood, Hanson, Shleifer, and Sørensen (2022) established the relevance of R-zones for financial crisis risk. One reason for the authors' focus on the postwar period was a constraint on the availability of household and business credit data. We also face such constraints. Nevertheless, pre-WWII data on household and business lending is available for *some* of the 18 countries that constitute our dataset.

Figure 8 shows very similar results when re-estimating the same model as above for the full sample period. At horizon $h = 2$, this increases the number of observations used in our estimation for household (business) sector R-zones from 945 (857) to 1012 (902). These additional data points do not have much impact on statistical precision and economic magnitudes of our estimation, a conclusion that also applies to a logistic model (right panels of Figure A9).

Figure 8: *The connection between loose monetary policy and R-zones: full-sample results.*

(a) *Household sector R-zone,* *$\overline{\text{stance}}$ lower by 100 bps*

(b) *Business sector R-zone,* *$\overline{\text{stance}}$ lower by 100 bps*

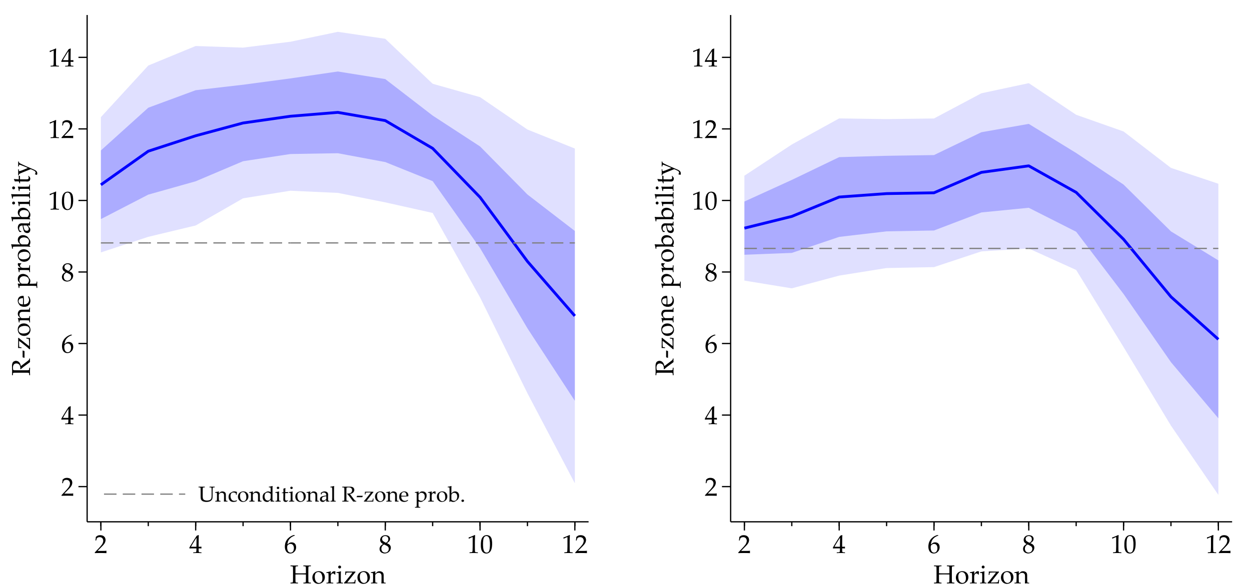


Notes: The same notes as in Figure 6 apply with one difference; all available data points of the JST Database are now used.

Figure 9: *The connection between loose monetary policy and "housing finance" R-zones.*

(a) *Post-WWII sub-sample,* *$\overline{\text{stance}}$ lower by 100 bps*

(b) *Full sample,* *$\overline{\text{stance}}$ lower by 100 bps*



Notes: R-zones are for housing sector only and using mortgage credit. The sample is restricted to the post-1949 period in panel (a). Apart from this modification, the same notes as in Figure 6 apply.

Since mortgage credit is closely related to household credit and allows us to exploit hundreds of additional pre-WWII data points, it is a natural extension to replace household credit with mortgage credit. We thus consider next the interaction between elevated house prices and mortgage credit with the goal to go further back in time. Motivated by earlier work that highlights the crucial role of the mortgage sector for financial stability considerations (Jordà, Schularick, and Taylor, 2015a), we refer to this interaction as a Red-zone too. Figure 9 establishes a relationship between loose monetary policy and the probability of what we now term *housing finance R-zones*. For comparison, we start again with the post-WWII period (panel (a)). The estimates shown in Figures 6 (a) and 9 (a) exhibit a similar pattern. The information in the stance of monetary policy is valuable for both post-WWII household sector and housing finance R-zones, and our measures are comparable in size. Panel (b) then shows results for the full sample that goes back, for some countries, to the 1870s. At horizon $h = 2$, we are now able to exploit 1368 observations in our estimation. They reveal that the predicted likelihood of entering a housing finance R-zone in five years is 10.2% (7.1%) when all covariates are at their sample mean but $\overline{stance}$ equals −1% (+1%).

Just like the positive relation between loose monetary policy stance and a heightened likelihood of financial crises, this link will also be confirmed in the following section within an LP-IV setting. To do so, we now leave LP-OLS estimates behind and spell out how we can employ an instrumental variable in this setting.

4 An instrumental variable approach

4.1 Construction of the instrument

Monetary policymakers do not play dice but set the stance of monetary policy for reasons. These reasons might go beyond the information set captured by our control variables.

One possible explanation for a loose stance of monetary policy is a series of expansionary and exogenous monetary policy shocks. In this section, we instrument $\overline{stance}$ with a series of such identified shocks. The strong first stage supports the relevance of monetary shocks for explaining policymakers' stance. The second stage invalidates claims that the positive link between loose monetary policy and elevated crisis and R-zone risk is simply due to endogeneity.

What are these shocks? We follow Jordà, Schularick, and Taylor (2020) who use the trilemma of international finance to construct a historical series of cross-country monetary policy shocks. The trilemma states that a country that pegs its currency to a base country and has an open capital account cannot conduct fully independent monetary policy at the same time (Obstfeld and Taylor, 2004). Rather, absence of international arbitrage implies that the pegging country has to adjust its policy rates, at least to some degree, in tandem with its base country, and the correlation will depend on the "hardness" of the peg. If the base country's exogenous interest rate changes do not take economic condition of the pegging country into account, then the pegging country's monetary policy response to them is exogenous.[14] In this case, the so-called trilemma IV is a valid instrument for interest rate changes of the pegging country. If variations in the trilemma IV are also of first-order relevance for shaping the stance of monetary policy, then it serves our purposes.

Formally, we define our instrumental variable (IV) as

$$z_{i,t} = \begin{cases} k_{i,t} \left(\Delta R_{b(i,t),t} - \Delta \hat{R}_{b(i,t),t} \right) & , \quad \text{if } \ q_{i,t} = 1\,; \\ 0 & , \quad \text{if } \ q_{i,t} = 0\,. \end{cases} \tag{7}$$

Here, $q_{i,t} \in \{0, 1\}$ is the exchange rate regime indicator; it equals 1 if i pegs its currency both in year t and $t-1$ and 0 else; $k \in [0, 1]$ denotes the re-scaled Quinn, Schindler, and Toyoda (2011) capital mobility indicator (1 if open).[15] $\Delta R_{b(i,t),t}$ is the nominal interest rate change in i's base country b in year t and $\Delta \hat{R}_{b(i,t),t}$ are corresponding predicted changes in $\Delta R_{b(i,t),t}$.[16] That is, our IV picks countries with well-established fixed exchange rate regimes, puts more weight on those countries with a high degree of capital mobility, and captures interest rate changes due to unpredictable interest rate movements in their respective base country. Note especially that pure interest parity, a correlation of 1, is *not required*; a valid IV just needs to have a positive correlation to satisfy the relevance condition, which ours does. See Jordà, Schularick, and Taylor (2020) for more details on this instrument and its construction.

An example of how trilemma-based identification works is provided by the now familiar tale of the eurozone in the 2000s. The instrument z as defined in equation (7) contains, for instance, non-zero values for the EMS and eurozone samples, with the exception of Germany which is treated as the base country for the economies of this currency union with base rates set by the Bundesbank and then the ECB.[17] Countries of the euro area exhibited heterogeneous growth rates in the 2000s. While the periphery was booming, the

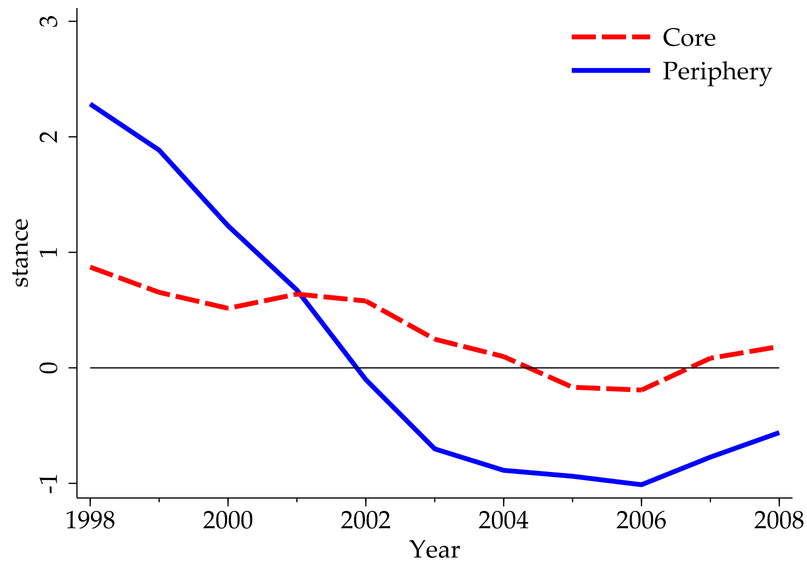
[14]See Jordà, Schularick, and Taylor (2015a) for a motivation of this identification assumption.

[15]The Quinn, Schindler, and Toyoda (2011) capital mobility indicator is not available for Ireland. The following IV estimates are therefore based on data for the remaining 17 countries of our sample. We checked that the OLS results presented in the other parts of this paper are not sensitive to the exclusion of Ireland.

[16]More precisely, $\Delta \hat{R}_{b(i,t),t}$ refers to predicted values from a cleaning regression, $\Delta R_{b(i,t),t} = \alpha_{b(i,t)} + \sum_{k=1}^{2} \beta_k \Delta R_{b(i,t),t-k} + \sum_{k=0}^{2} \mathbf{\Gamma}_k \mathbf{X}_{b(i,t),t-k} + e_{b(i,t),t}$. Here, $\mathbf{X}$ includes the inflation rate and changes in log real GDP p.c., log real stock prices, log real wages, inflation, and the credit-to-GDP ratio.

[17]For a complete overview of the assignment of base countries to pegging countries, see Jordà, Schularick, and Taylor (2020, Table 1).

Figure 10: *The stance of monetary policy in the eurozone before the Global Financial Crisis.*



Notes: The figure shows the unweighted average of $\overline{stance}$ as defined in equation (4) for the core countries (Belgium, Denmark, France, Germany, Netherlands) and for the periphery countries (Ireland, Italy, Portugal, Spain) of the eurozone.

core grew only moderately. If monetary policy reacted to economic conditions in the core only, then interest rate decisions of the European Central Bank were too accommodative for countries such as Spain and Ireland. Over time, this should imply a looser stance of monetary policy in the periphery vis-à-vis the core. Indeed, the literature argues that in the years before the Global Financial Crisis, monetary policy of the European Central Bank was in line with macroeconomic conditions in the core countries but too accommodative for the periphery (Nechio, 2011; Lothian, 2014).

Does our stance measure match up with these judgements from independent analyses? Yes, which provides additional confidence from a specific historical episode. Our estimates of the stance of eurozone monetary policy reflect this as illustrated in Figure 10. For the periphery countries of the eurozone, our stance measure features a sharp downward trend and is *too loose* over the years that preceded the crisis. The fall of the stance in the core, on the other hand, was only moderate. For these core countries, our estimates do not indicate a significant deviation from a neutral stance of monetary policy.

We are interested in the variation in the stance of monetary policy induced by series of accommodative monetary policy shocks such as the one that hit the European periphery in the 2000s.[18] To this end, we instrument $\overline{stance}$ with 10 lags of $z_{i,t}$. Recall that we define

[18] In the model of Boissay, Collard, Galí, and Manea (2022), it is precisely such a long period of expansionary monetary policy shocks that predates the average financial crisis.

the stance of monetary policy as the moving-averaged difference between real rate and natural rate from year $t-4$ to t (see equation (4)). By construction, a monetary policy shock in year $t-1$ can only have a limited effect on $\overline{stance}$. In contrast, a series of medium-term expansionary policy shocks should have a considerable effect on a central bank's stance.

This intuition behind our IV choice is confirmed when we look at Table A1 which presents the first stage regression. A history of medium-term contractionary monetary policy shocks is linked to a significantly tighter stance of monetary policy and vice versa. Together, the instrumental variables present a strong first stage as indicated by the Kleibergen-Paap (2006) test for weak instruments.[19]

4.2 Financial crises

When we instrument for $\overline{stance}$ with our trilemma IV, our core result—a loose stance is associated with heightened financial crisis risk in the medium term—remains in place.

Indeed, we find that the impulse responses here exhibit a very similar pattern. This is shown in Figure 11. Here, the instrumented stance predicts financial crises starting at a horizon of around five years at statistically significant levels. As in the previous LP-OLS setup (Figure 3), bootstrap confidence bands confirm the significance of the estimates, as shown by the shaded areas. Furthermore, similar results are obtained when focusing on the post-WWII period (A10), and the relationship between loose monetary policy and the *exact* starting year of systemic banking crises in the medium term is also significant at the 5% level (Figure A11).

Point estimates are now larger, likely because IV mitigates attenuation bias under OLS due to measurement error coming from r^*, as noted. In our baseline specification without time fixed effects, the likelihood of entering a financial crisis 5 to 7 years ahead is higher by 5.5 pps when our measure for the stance of monetary policy is looser by 1 pp. As shown in panel (a) of Figure 11, this implies a crisis probability of 16% when $\overline{stance} = -1\%$. At higher horizons, the medium-term crisis risk more than doubles from 10.5% to over 20%.

4.3 Credit and asset market overheating

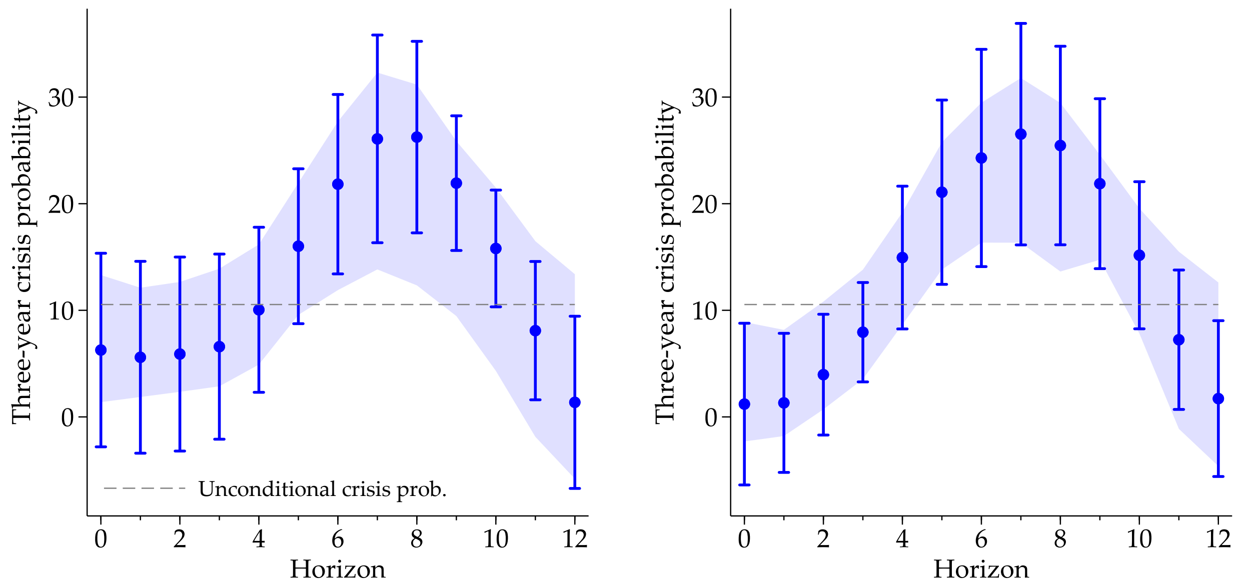
In Figure 10, we showed the drifting apart of the stance of eurozone monetary policy between core and periphery in the 2000s. We argued this was due to a central bank that

[19] The same control variables we used in the previous sections enter the first-stage regression. Notice that this includes lags of the dependent variable under study. In the following, we present IV estimates for different dependent variables. Hence, the corresponding first stages slightly differ. For the sake of brevity, Table A1 only reports the first-stage regression with lags of our financial crisis variable in the vector of control variables. We checked, though, that the Kleibergen-Paap (2006) test for weak instruments yields a statistic above 10 in all other first-stage regressions as well.

Figure 11: *The effect of loose monetary policy on crisis risk: second stage.*

(a) *Baseline, $\overline{stance}$ lower by 100 bps*

(b) *Decade fixed effects, $\overline{stance}$ lower by 100 bps*



Notes: We re-estimate model (5) by 2SLS. In panel (a), the same control variables as before, outlined in section 2.2, are used. In panel (b), we replace the global control variables by decade fixed effects. $\overline{stance}_{i,t}$ is instrumented with $\{z_{i,t-k}\}_{k=1}^{10}$ as defined in equation (7). The corresponding first stages are presented in columns (1) and (2) of Table A1. The points show IV estimates of $\{-100\beta^h\}_{h=0}^{12}$. Apart from these modifications, the same notes as in Figure 3 apply.

focused on the moderately growing core while ignoring the booming periphery. This line of reasoning motivates the choice of our set of instrumental variables. A series of interest rate cuts in the base country that is too loose given macroeconomic conditions in the pegging country can push interest rates of the pegging country below neutral.

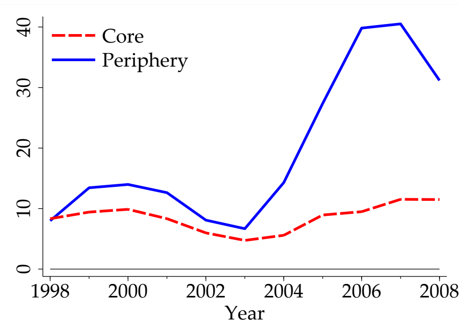
Does this arguably exogenous variation in the stance of monetary policy raise the likelihood of entering R-zones, that is periods of credit and asset market overheating that have been shown to be of first-order relevance for financial stability considerations?

Figure 12 provides the first evidence for that. The divergence of the eurozone countries' stance of monetary policy is mirrored in those financial variables we are interested in. While core and periphery experienced largely synchronized pre-2008 stock price dynamics, the credit and house price boom in the 2000s was far more pronounced in the periphery than in the core, i.e., in those countries that faced too loose a stance of monetary policy.

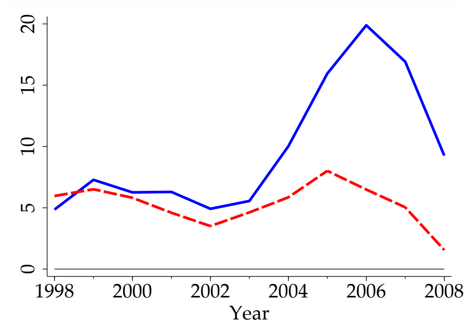
So much for that episode, but is this a general result in our data? Yes. Figure 13 generalizes this indicative evidence with impulse responses from LP-IV estimation. These estimates confirm the LP-OLS estimates reported in Figure 6; a loose stance of monetary

Figure 12: *Credit and asset prices in the eurozone before the Global Financial Crisis.*

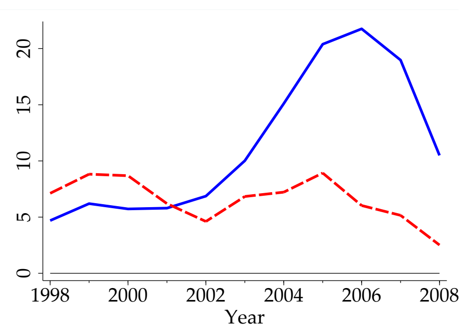
(a) *Total credit*



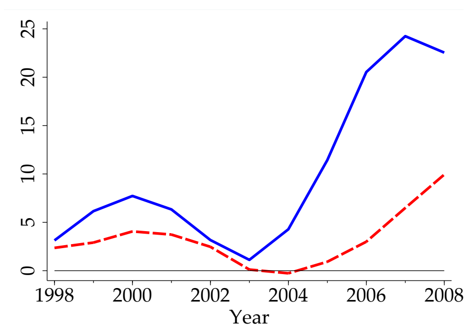
(b) *Household credit*



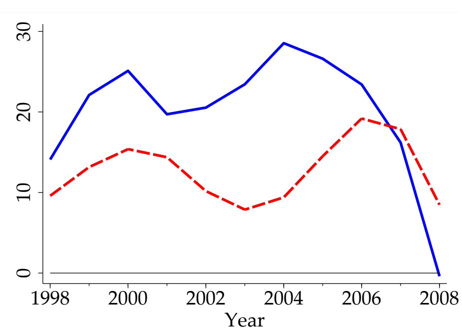
(c) *Mortgage credit*



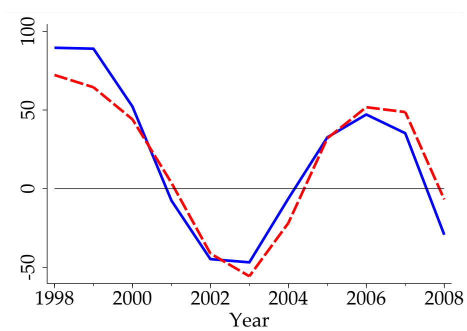
(d) *Business credit*



(e) *House prices*



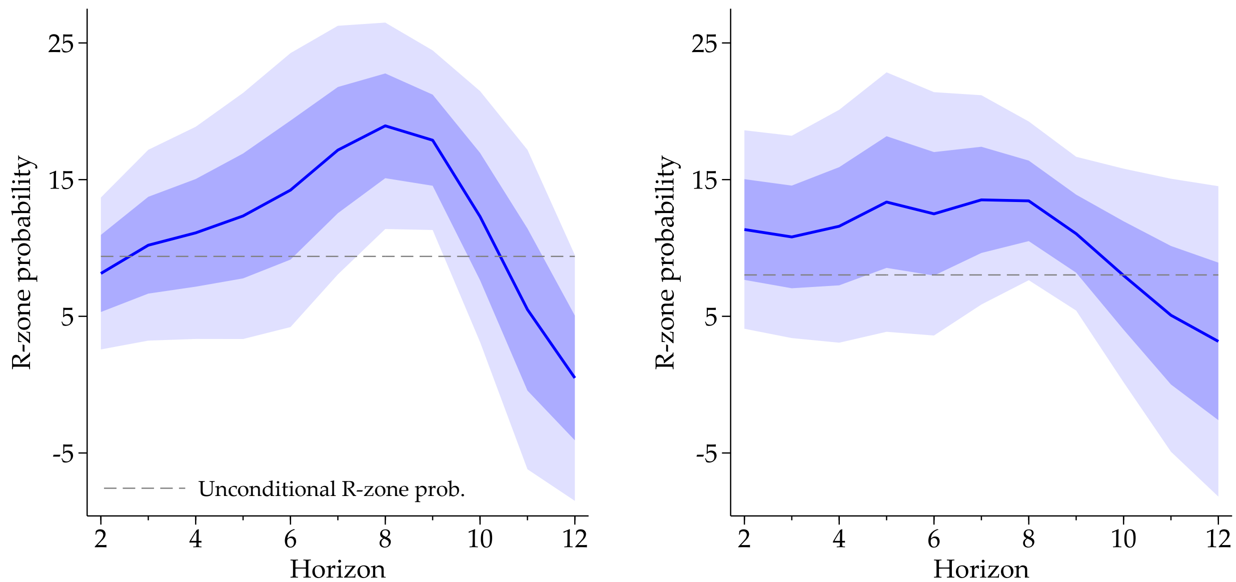
(f) *Stock prices*



Notes: Motivated by the definition of R-zones (equation (6)), the figure shows unweighted averages of $\Delta_3\,(100 Debt/GDP)$ (first four panels) and $\Delta_3(100 \log Price)$ (last two panels) for the core countries (Belgium, Denmark, France, Germany, Netherlands) and for the periphery countries (Ireland, Italy, Portugal, Spain) of the eurozone. The type of *Debt* and *Price* is specified in the titles of the panels.

Figure 13: *The effect of loose monetary policy on post-WWII R-zones: second stage.*

(a) *Household sector R-zone, $\overline{stance}$ lower by 100 bps* (b) *Business sector R-zone, $\overline{stance}$ lower by 100 bps*



Notes: The same notes as in Figure 6 apply with one difference. The figure here shows IV estimates and corresponding 95% (light) and 68% (dark) confidence intervals based on country-based cluster-robust standard errors. $\overline{stance}_{i,t}$ is instrumented with $\{z_{i,t-k}\}_{k=1}^{10}$ as defined in equation (7).

policy raises the likelihood of credit and asset market overheating as defined by Greenwood, Hanson, Shleifer, and Sørensen (2022) in the following years at economically relevant levels in the post-WWII period. The same is true for housing finance R-zones which we introduced in the previous section (Figure A12) and in the full sample (Figure A13). As already mentioned earlier in the context of the OLS estimates, the larger degree of statistical uncertainty in panel (b) of Figure 13 might reflect heterogeneity in corporate credit (Müller and Verner, 2023).

5 A growth-risk trade-off

Finally, we note that a loose stance of monetary policy also has potential benefits, as well as costs. Loose financial conditions and increased risk taking may not be a bad thing *per se*. They might, for example, enhance consumption smoothing by relaxing financial constraints or raise innovation and efficiency by providing more investment capital.

However, our historical evidence suggests that running such a high-pressure economy may not be sustainable in general. In the following, we argue that potential short-term gains come at the considerable cost in the form of heightened risk of disasters in real economic activity. As outlined in the Introduction, this closes the circle between the conduct of monetary policy, financial fragility, and real activity, while also speaking to the Growth-at-Risk literature (Adrian, Boyarchenko, and Giannone, 2019; Adrian, Grinberg, Liang, Malik, and Yu, 2022).

We have shown before that too accommodative a stance of monetary policy breeds financial instability in the medium term. We now complete this picture of heightened systemic risk by looking at the consequences for real growth. To do so, we remain within the framework of the previous sections and define an indicator that captures periods of exceptionally low real growth from a historical perspective. More precisely, we define Y as real GDP per capita, and construct a 3-year low-growth indicator as

$$\textit{Low-Output-Growth}_{i,t} = \mathbb{1}\{\Delta_3 \left(\log Y_{i,t}\right) < 20^{th}\, \textit{percentile}\},$$

and return once more to model (5). Now, $B_{i,t} = \textit{Low-Output-Growth}_{i,t}$. As in the previous parts of this paper, the vector of control variables $\mathbf{X}$ remains unchanged.

Panel (a) of Figure 14 shows estimates of such a model. Consistent with the non-existent association between a loose stance and elevated short-term crisis risk (Figure 3), the likelihood of entering the left tail of the cross-country real growth distribution stays put in the immediate aftermath of accommodative monetary policy. In the short term, only the R-zone indicators show a positive response as shown in the previous sections. This short period of tranquility is followed by a rise in the likelihood of sharp downturns in real activity.

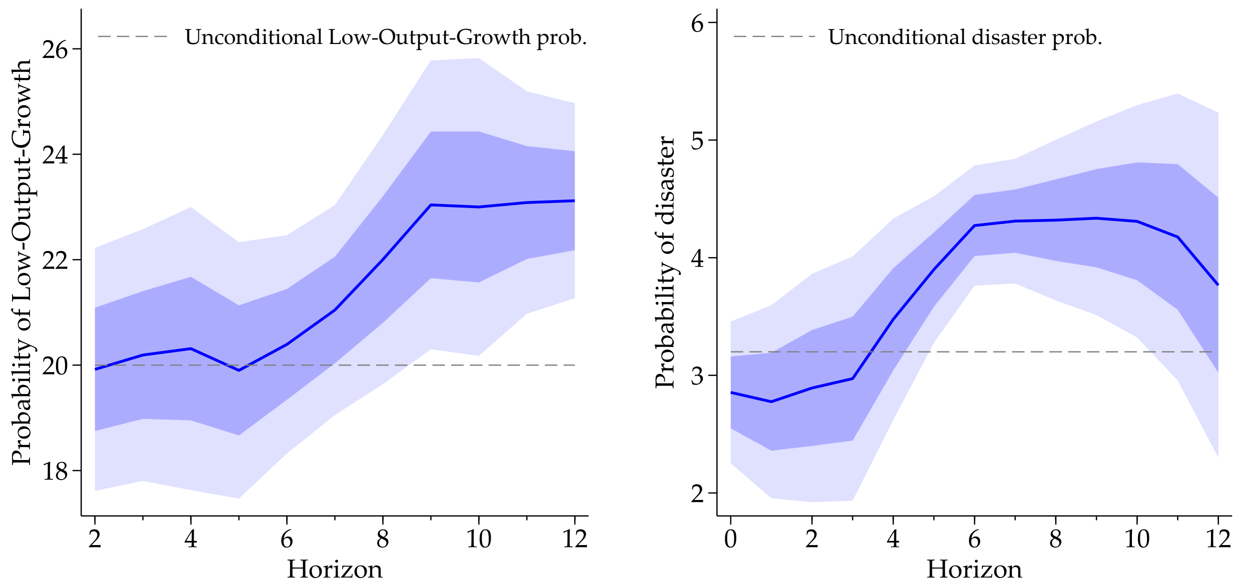
The connection between $\overline{\textit{stance}}$ and the *Low-Output-Growth* indicator starts to take shape at a horizon of 6 years. That is, when monetary policy is more accommodative, the likelihood of historically low GDP growth rates 4 to 6 years in advance is higher. This relation becomes significant at the 32% (5%) level at a horizon of 7 (9) years and remains so in subsequent years. When our stance variable is looser by 1 pp, the likelihood that the *Low-Output-Growth* indicator turns on 7 (9) years ahead is 1 (3) pps higher. This, in turn, implies historically low growth rates from $h = 5$ to $h = 9$.

We cannot really interpret the 20th percentile of the pooled cross-country 3-year real growth distribution as a *disaster*. The growth rate at this position in the distribution is still positive, and is 1.32% in our sample. Of course, a widely-used definition of actual economic disasters was put forward by Barro and Ursúa (2008). They define such events as periods of peak-to-trough falls in real GDP per capita of at least 10%. Historically, such disasters in real economic activity have occurred at the highest frequency during (or around) the two world wars, with quite a few more during the Great Depression years of the early 1930s (Barro and Ursúa, 2008).

Figure 14: *The connection between loose monetary policy and the left tail of real growth.*

(a) *Low-Output-Growth indicator, $\overline{\textit{stance}}$ lower by 100 bps*

(b) *Barro and Ursúa (2008) disasters, $\overline{\textit{stance}}$ lower by 100 bps*



Notes: The figure shows estimates of $-100\beta^h$ of model (5) for different horizons h relative to the unconditional full-sample probability that $B_{i,t} = 1$. In panel (a), $B_{i,t}$ = *Low-Output-Growth* as defined in the text. In panel (b), $B_{i,t} = 1$ if country i experiences a real disaster in year t and 0 else. As in Barro and Ursúa (2008), a real disaster is defined as a peak-to-trough fall in real GDP per capita of at least 10%; peaks and troughs of real GDP per capita are defined using the Bry-Boschan (1971) algorithm. For both panels, $b_{i,t}$ is empty, since the vector of control variables $\mathbf{X}$ already contains past information about real GDP p.c. growth. Shaded areas indicate 95% (light) and 68% (dark) confidence intervals based on Driscoll-Kraay (1998) standard errors with *ceiling*$(1.5 \times h)$ lags.

It turns out that 3.2% of the observations in our sample—which excludes the world wars as described in section 2—are characterized by such disasters. This definition is therefore closer to the concept of Growth-at-Risk that is usually based on the 5th percentile of the GDP growth distribution (e.g., Aikman, Bridges, Hoke, O'Neill, and Raja, 2019; Franta and Gambacorta, 2020; Lloyd, Manuel, and Panchev, 2023; Adrian, Grinberg, Liang, Malik, and Yu, 2022). The above-stated finding that a loose stance of monetary policy predicts left-tail events in medium-term real growth remains in place when substituting the *Low-Output-Growth* indicator with Barro and Ursúa (2008) disaster events. This is shown in panel (b) of Figure 14 and this result even points towards slightly *lower* disaster risk in the short term when the policy stance is loose, offset by much higher risk later on.

6 Conclusion

This study provides the first evidence that the stance of monetary policy has implications for the stability of the financial system. Whilst our LP-OLS estimates suggest a clear predictive relationship, our LP-IV estimates provide more robust evidence for a causal interpretation. A loose stance over an extended period of time leads to increased financial fragility several years down the line. The source of this fragility is associated with swings in those financial variables that have been identified by the literature as harbingers of financial turmoil.

Policymakers should take seriously the possible dangers that can be ignited by keeping policy rates lower for longer, and thus weigh the potential short-run gains of loose monetary policy against the substantial risks of extremely adverse medium-term crisis consequences. Such policies increase the likelihood of financial crises and thus the risk of high social, political, as well as economic costs as time goes by.

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A1 Figures and Tables

Table A1: *First stage.*

Dep. var.: $\overline{stance}_{i,t}$	Full sample (1)	(2)	Post-WW2 (3)	(4)
$z_{i,t-1}$	0.063	-0.033	0.119	-0.013
	(0.073)	(0.063)	(0.094)	(0.077)
$z_{i,t-2}$	0.218***	0.174***	0.301***	0.210***
	(0.054)	(0.058)	(0.084)	(0.080)
$z_{i,t-3}$	0.263***	0.276***	0.406***	0.368***
	(0.052)	(0.054)	(0.086)	(0.088)
$z_{i,t-4}$	0.326***	0.340***	0.413***	0.397***
	(0.049)	(0.048)	(0.074)	(0.082)
$z_{i,t-5}$	0.235***	0.295***	0.386***	0.376***
	(0.035)	(0.034)	(0.076)	(0.059)
$z_{i,t-6}$	0.180***	0.279***	0.305***	0.359***
	(0.042)	(0.038)	(0.087)	(0.069)
$z_{i,t-7}$	0.164***	0.186***	0.289***	0.234***
	(0.048)	(0.042)	(0.100)	(0.073)
$z_{i,t-8}$	0.155***	0.237***	0.267***	0.307***
	(0.050)	(0.055)	(0.100)	(0.089)
$z_{i,t-9}$	0.111**	0.199***	0.149**	0.266***
	(0.045)	(0.043)	(0.074)	(0.072)
$z_{i,t-10}$	0.082*	0.145***	0.101*	0.184***
	(0.046)	(0.049)	(0.060)	(0.066)
Controls	✓	✓	✓	✓
Time FEs	✗	Decade	✗	Decade
KP weak IV	47.16	27.95	15.05	29.17
Observations	1297	1297	959	959

Notes: We re-estimate model (5) by two-stage least squares. $\overline{stance}_{i,t}$ is instrumented with $\{z_{i,t-k}\}_{k=1}^{10}$ as defined in equation (7). The table shows first-stage estimates of $\overline{stance}_{i,t}$ on these instruments. Control variables are the same as before and outlined in section 2.2. In columns (2) and (4), we replace the global control variables by decade fixed effects. Country-based cluster-robust standard errors are shown in parentheses. *, **, and *** denote significance at the 10%, 5%, and 1% level, respectively. *KP weak IV* refers to the Kleibergen-Paap (2006) Wald rk F-statistic.

Figure A1: *By-country time series of* $r^*_{i,t} = \bar{r}^w_t + \bar{r}^i_t$ *and* $\overline{stance}_{i,t}$.

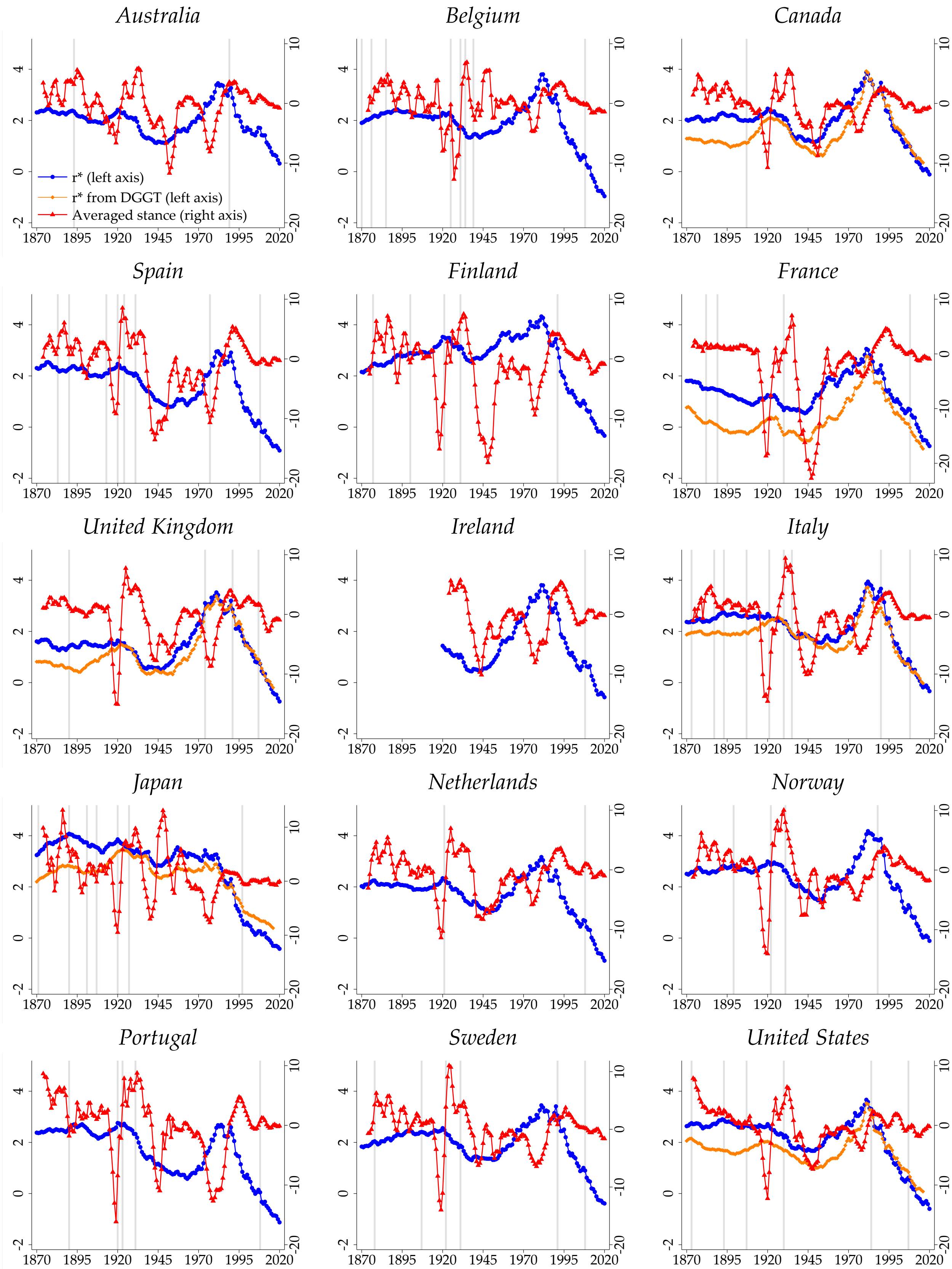
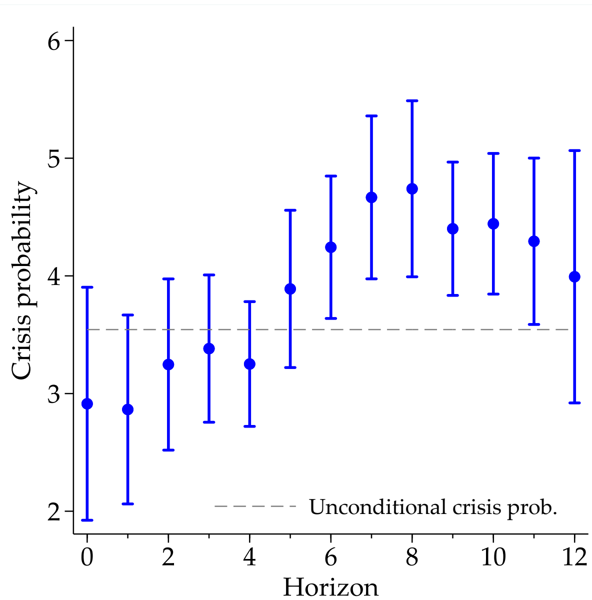
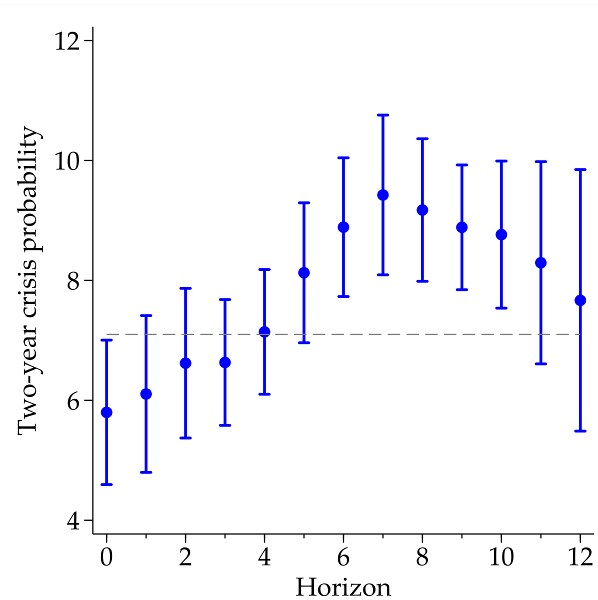


Figure A2: *The connection between loose monetary policy and financial crises: different horizons of the dependent variable.*

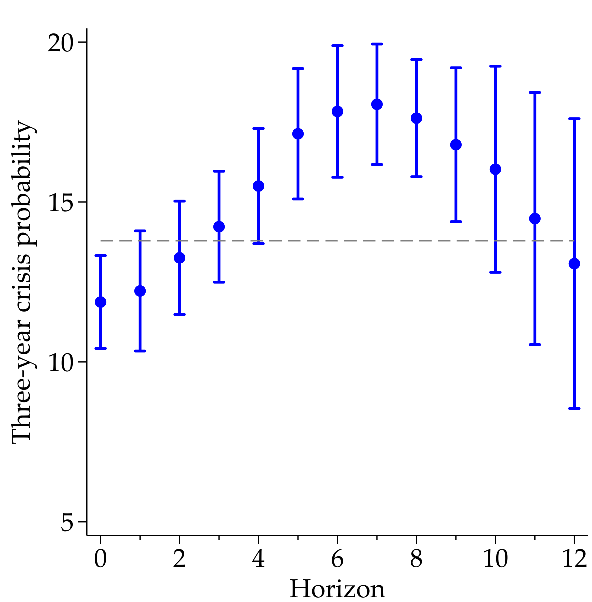
(a) $F = 0$, $\overline{stance}$ *lower by 100 bps*



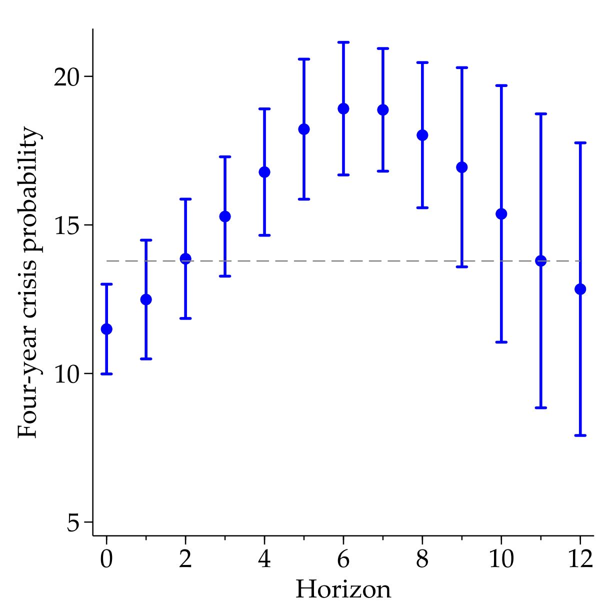
(b) $F = 1$, $\overline{stance}$ *lower by 100 bps*



(c) $F = 3$, $\overline{stance}$ *lower by 100 bps*

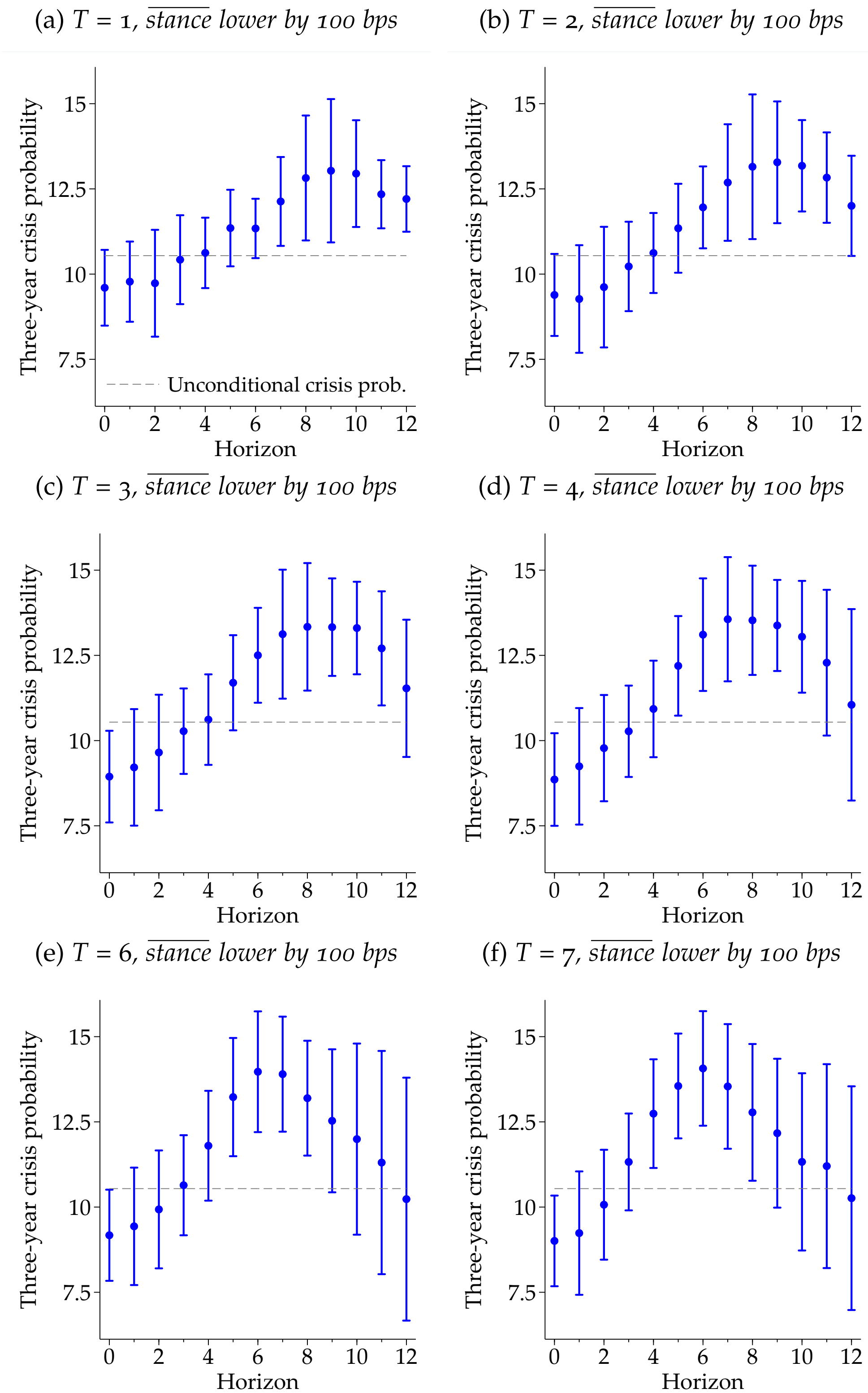


(d) $F = 4$, $\overline{stance}$ *lower by 100 bps*



Notes: We re-estimate model (5) with the dependent variable defined as $B_{i,t} = \max\{crisis_{i,t}, crisis_{i,t+1}, \ldots, crisis_{i,t+F}\}$. The figure shows estimated coefficients of $\{-100\beta^h\}_{h=0}^{12}$ for different values of F relative to the unconditional full-sample (F+1)-year crisis probability. Apart from this modification, the same notes as in Figure 3 apply.

Figure A3: *The connection between loose monetary policy and financial crises: averaging $r - r^*$ over a different number of years.*

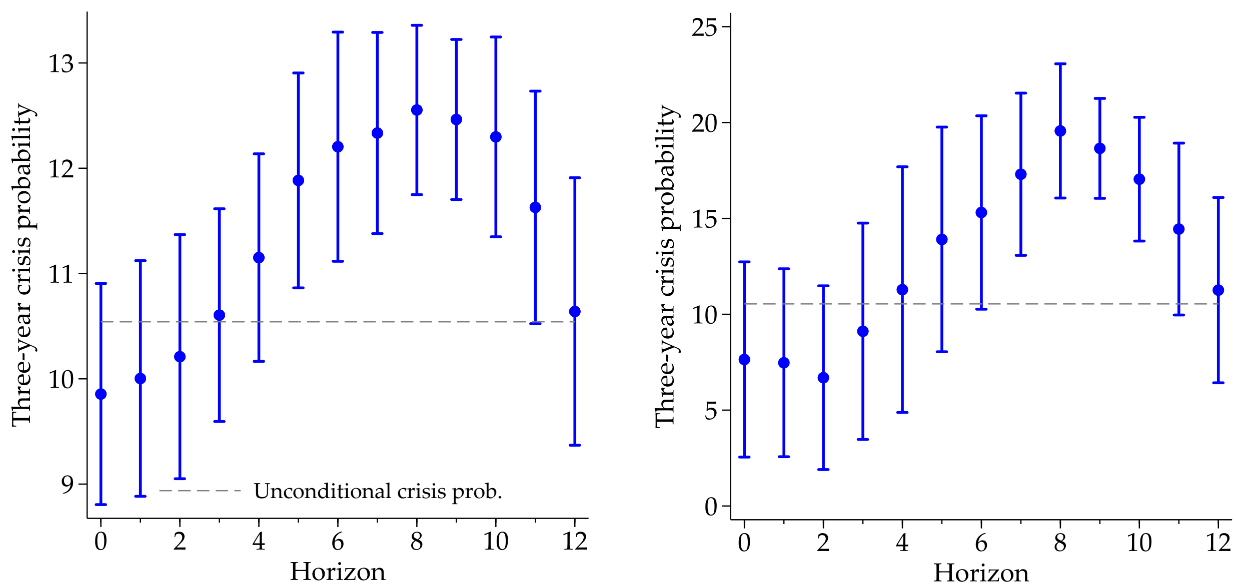


Notes: The independent variable of interest is now defined as $\frac{1}{T}\sum_{k=0}^{T-1}(r_{i,t-k} - r^*_{i,t-k})$. Apart from this modification, the same notes as in Figure 3 apply.

Figure A4: *Estimated financial crisis risk within a logistic model.*

(a) $\overline{stance}$ *lower by 100 bps*

(b) $\mathbb{1}\{\overline{stance} < 20^{th}$ *percentile*$\} = 1$

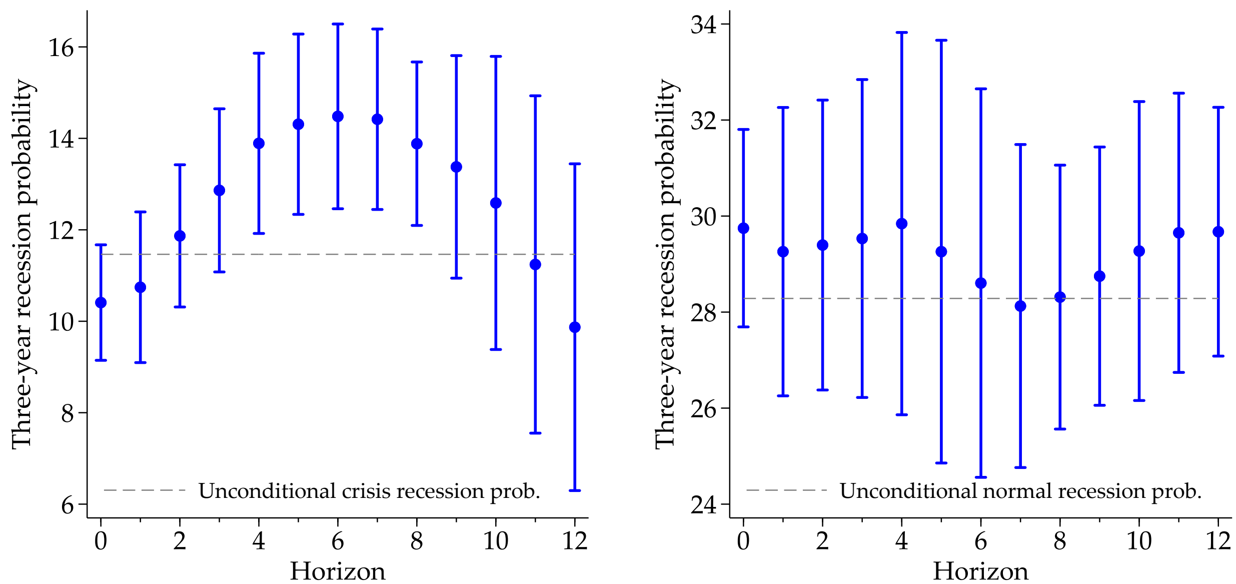


Notes: We take again model (5), replace the right-hand-side of the equation by $\log\left(\frac{p_{i,t}}{1-p_{i,t}}\right)$ with $p_{i,t} = Pr(B_{i,t} = 1)$, and estimate the model by maximum likelihood. Point estimates refer to marginal effects of $\overline{stance}$ evaluated at the sample means of the covariates relative to the unconditional full-sample three-year crisis probability. Control variables are the same as before and outlined in section 2.2. Bars indicate 95% confidence intervals based on robust standard errors. As in Figure 3, positive estimates suggest a positive relation between a loose stance of monetary policy and crisis risk.

Figure A5: *Normal and financial crisis recessions.*

(a) *Financial crisis recessions, $\overline{\text{stance}}$ lower by 100 bps*

(b) *Normal recessions, $\overline{\text{stance}}$ lower by 100 bps*

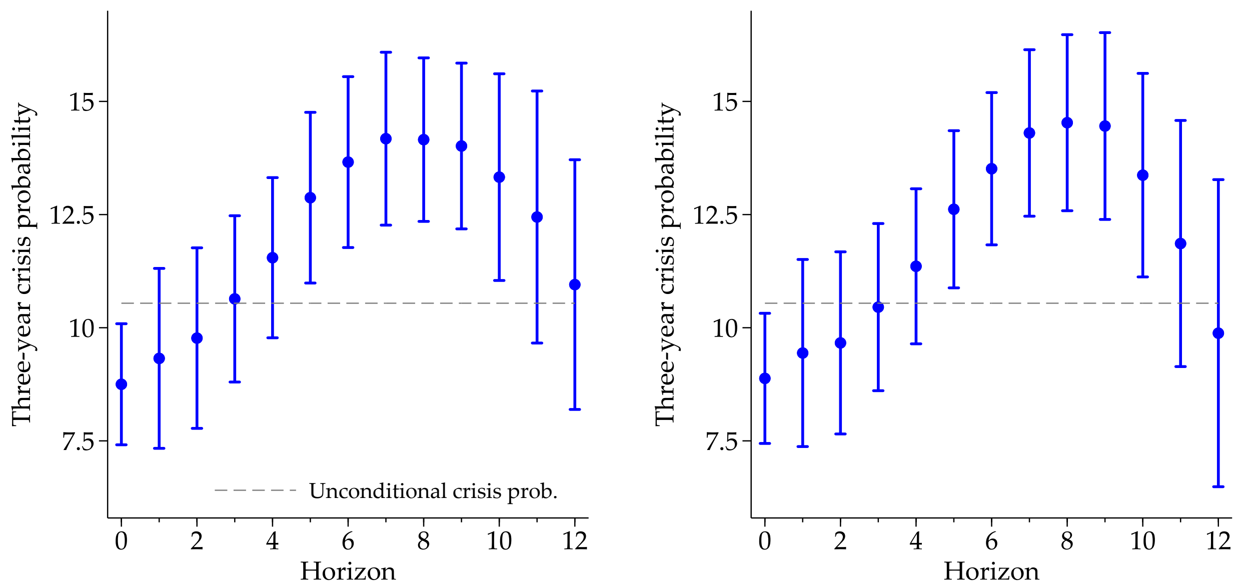


Notes: The same notes as in Figure 3 apply with one difference; the financial crisis event indicator is replaced by a financial crisis recession event indicator (panel (a)) or a normal recession event indicator (panel (b)). To construct these indicators, we identify peaks and troughs of real GDP p.c. using the Bry-Boschan (1971) algorithm. As in Jordà, Schularick, and Taylor (2016), we then define the peak dates as the starting years of recessions and group these recessions into *normal recessions* and *financial crisis recessions*. Financial crisis recessions are those recessions that occur within a ±2-year window of financial crises. Normal recessions are the remaining ones.

Figure A6: *The connection between loose monetary policy and financial crises: including decade fixed effects.*

(a) *Global variables and decade FEs,* $\overline{\textit{stance}}$ *lower by 100 bps*

(b) *Only decade FEs,* $\overline{\textit{stance}}$ *lower by 100 bps*

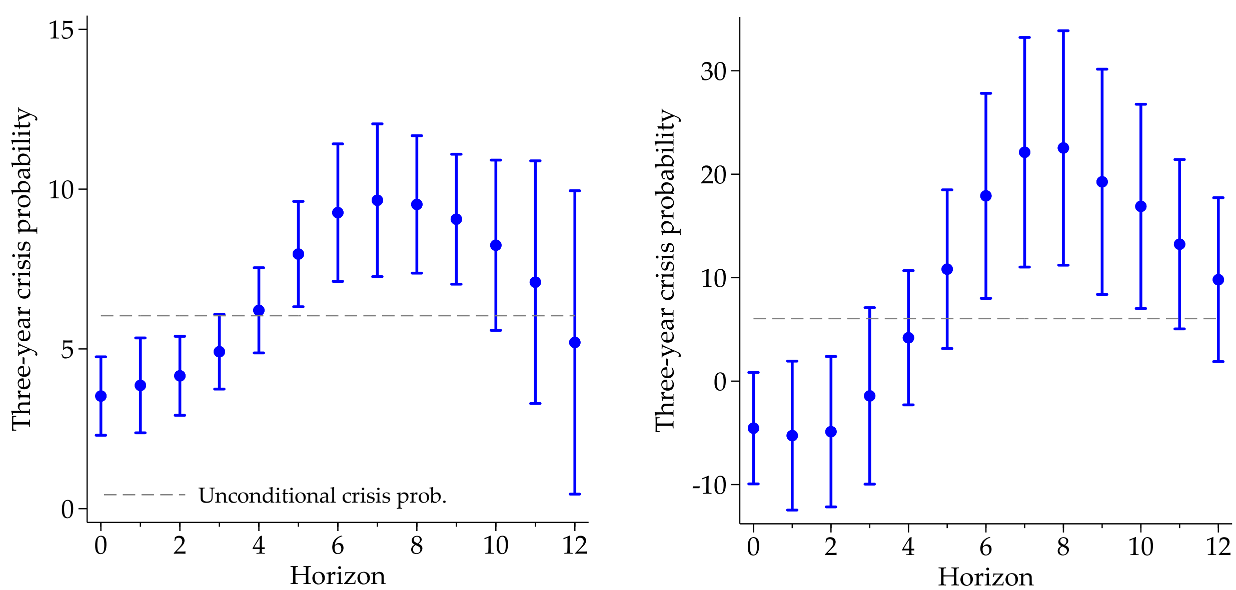


Notes: The same notes as in Figure 3 apply with one difference; we enrich the set of controls with decade fixed effects (panel (a)) or replace the global variables with decade fixed effects (panel (b)).

Figure A7: *The connection between loose monetary policy and post-WWII financial crises.*

(a) $\overline{\textit{stance}}$ *lower by 100 bps*

(b) $\mathbb{1}\{\overline{\textit{stance}} < 20^{th}\ \textit{percentile}\} = 1$

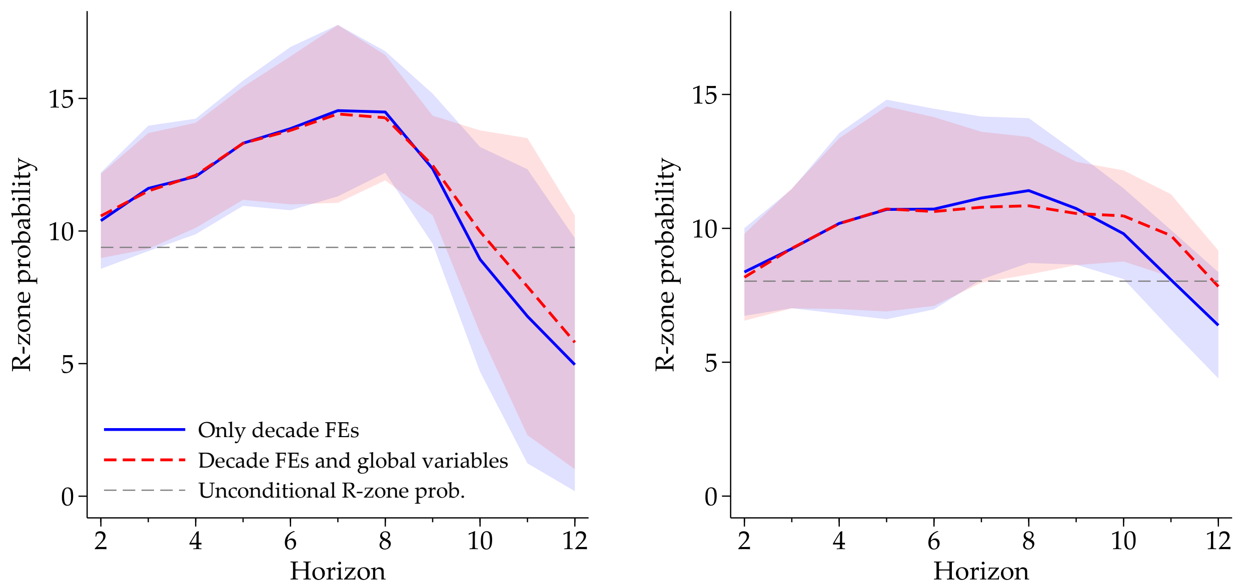


Notes: The same notes as in Figure 3 apply with one difference; we use only data for the post-1945 period.

Figure A8: *The connection between loose monetary policy and post-WWII R-zones: including decade fixed effects.*

(a) *Household sector, $\overline{\text{stance}}$ lower by 100 bps*

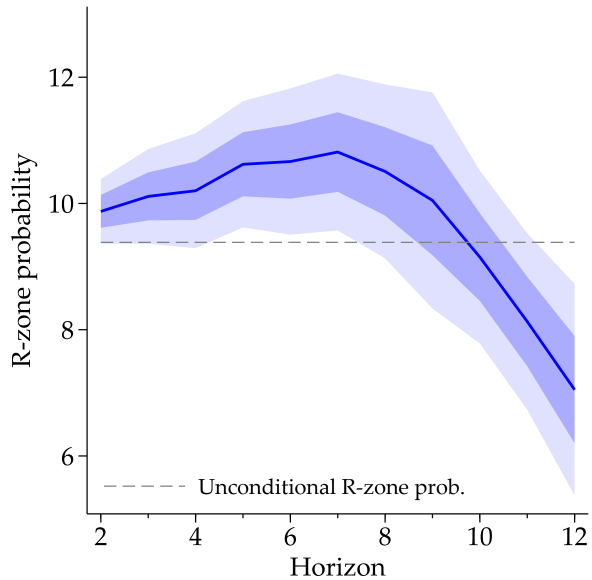
(b) *Business sector, $\overline{\text{stance}}$ lower by 100 bps*



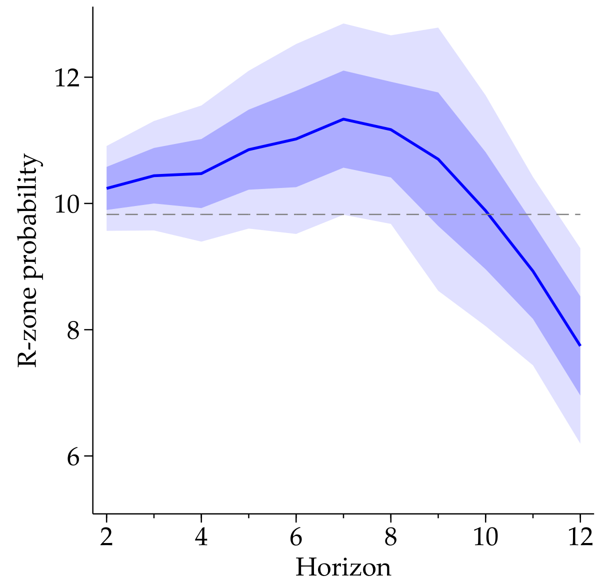
Notes: The same notes as in Figure 6 apply with one difference; decade fixed effects are added to the model, either in place of global control variables (solid blue lines) or in addition to global control variables (dotted red lines). Shaded areas indicate 95% confidence intervals based on Driscoll-Kraay (1998) standard errors with *ceiling*(1.5 × *h*) lags.

Figure A9: *Estimated R-zone risk within a logistic model.*

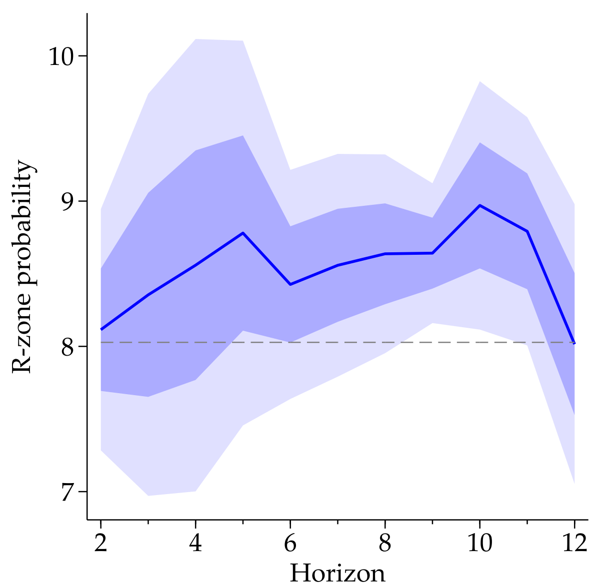
(a) *Household sector, post-WWII sub-sample,*
$\overline{\text{stance}}$ lower by 100 bps



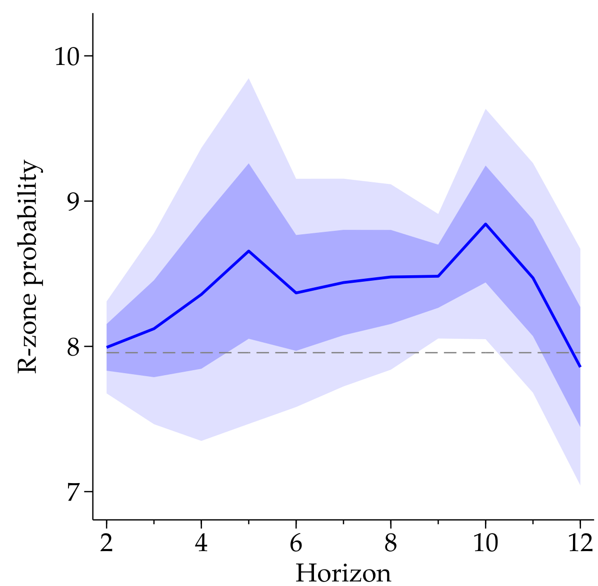
(b) *Household sector, full sample,*
$\overline{\text{stance}}$ lower by 100 bps



(c) *Business sector, post-WWII sub-sample,*
$\overline{\text{stance}}$ lower by 100 bps



(d) *Business sector, full sample,*
$\overline{\text{stance}}$ lower by 100 bps

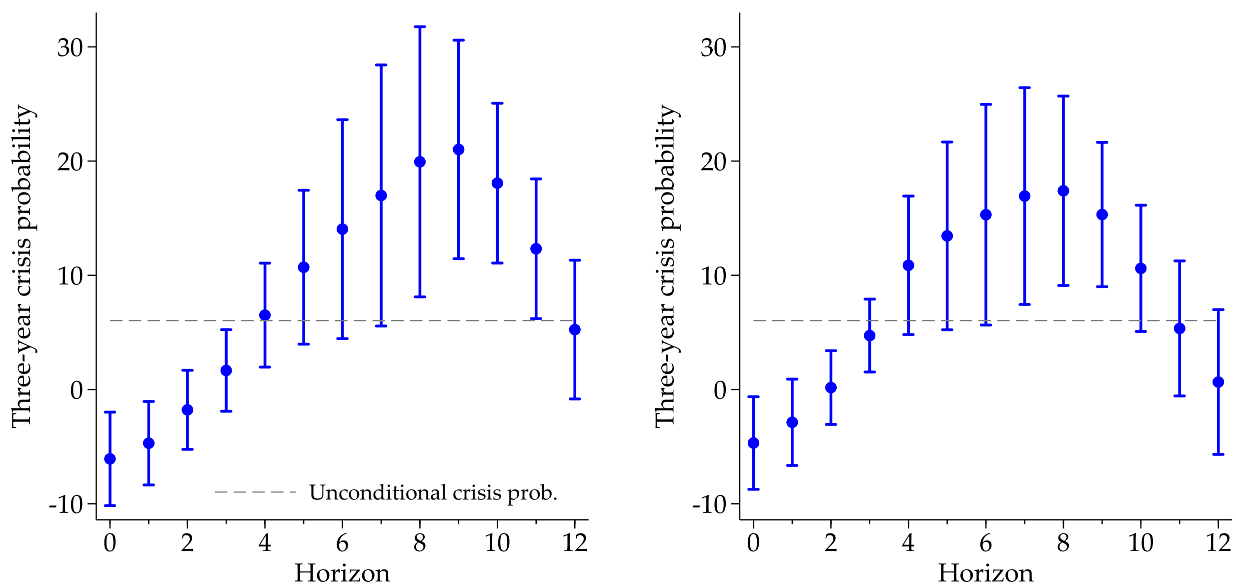


Notes: Similar to the financial crisis part (Figure A4), the figure shows the estimated R-zone risk within a logistic model.

Figure A10: *The effect of loose monetary policy on crisis risk: second stage, post-WWII period.*

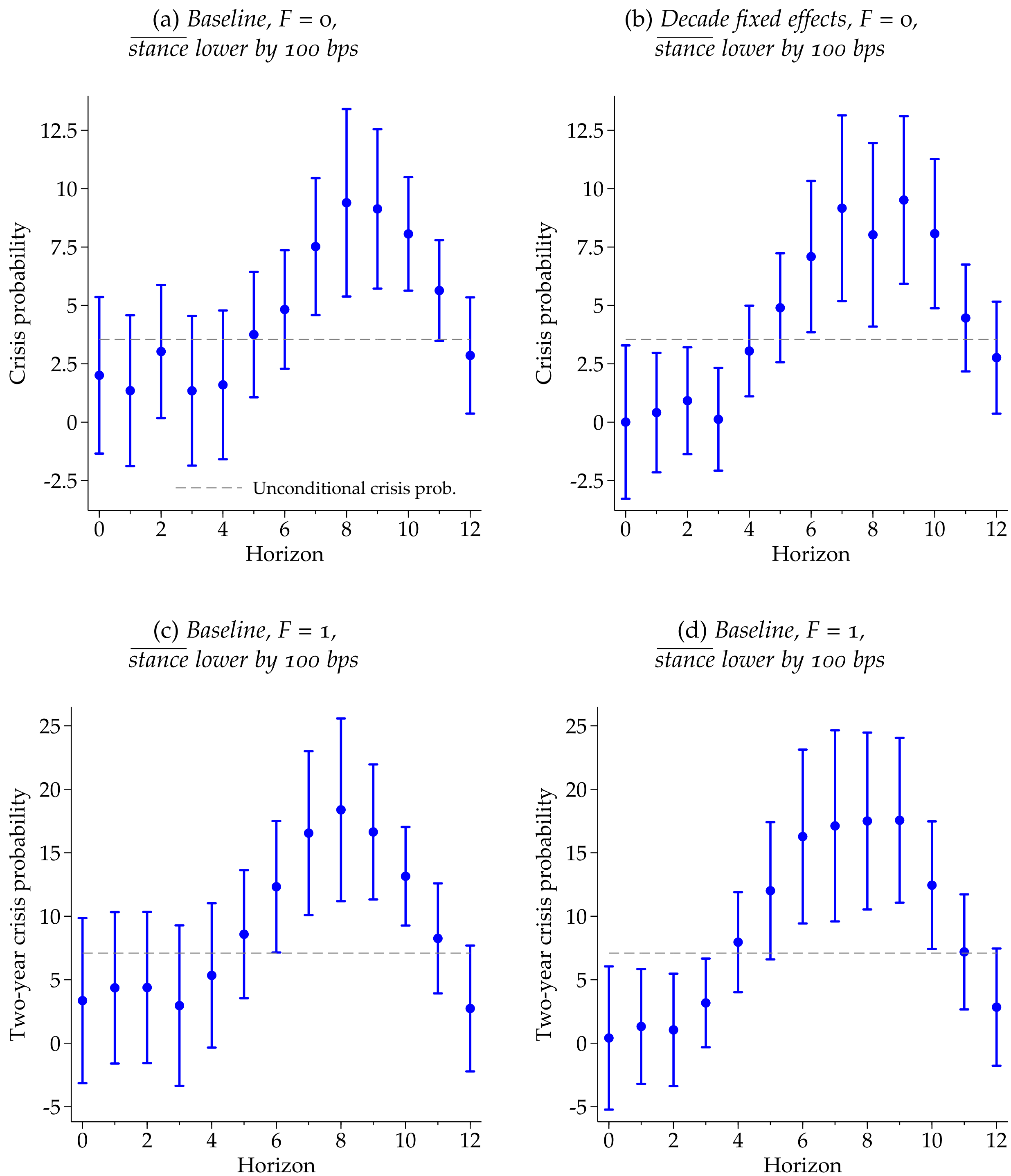
(a) *Baseline, $\overline{\text{stance}}$ lower by 100 bps*

(b) *Decade fixed effects, $\overline{\text{stance}}$ lower by 100 bps*



Notes: The same notes as in Figure 11 apply with one difference; we use only data for the post-1945 period. The corresponding first stages are presented in columns (3) and (4) of Table A1.

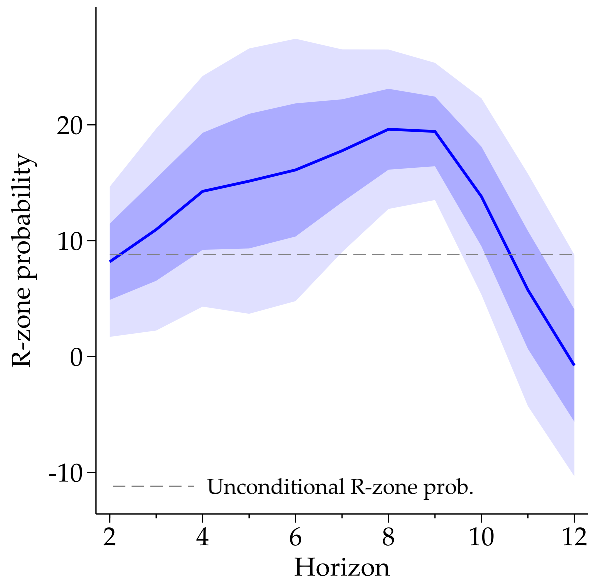
Figure A11: *The effect of loose monetary policy on crisis risk: second stage, different horizons of the dependent variable.*



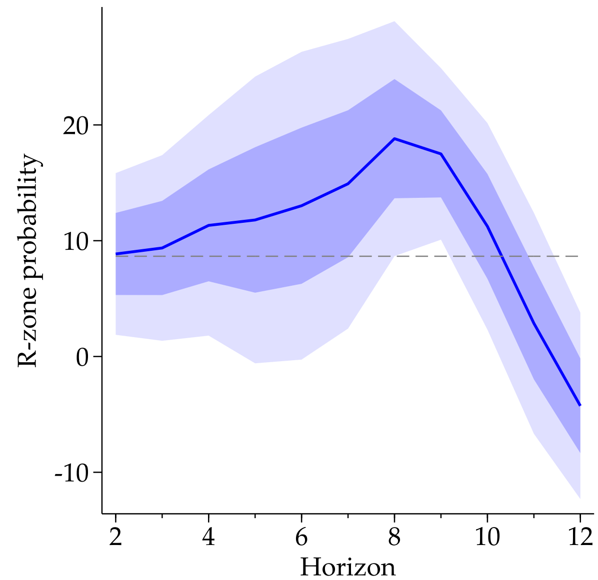
Notes: In panels (a) and (b), $B_{i,t} = crisis_{i,t}$. In panels (c) and (d), $B_{i,t} = \max\{crisis_{i,t}, crisis_{i,t+1}\}$. The figure shows estimated coefficients of $\{-100\beta^h\}_{h=0}^{12}$ relative to the unconditional full-sample one-year (upper panels) or two-year (lower panels) crisis probability. Apart from this modification, the same notes as in Figure 11 apply.

Figure A12: *The effect of loose monetary policy on "housing finance" R-zones: second stage.*

(a) *Post-WWII sub-sample,*
$\overline{\text{stance}}$ lower by 100 bps



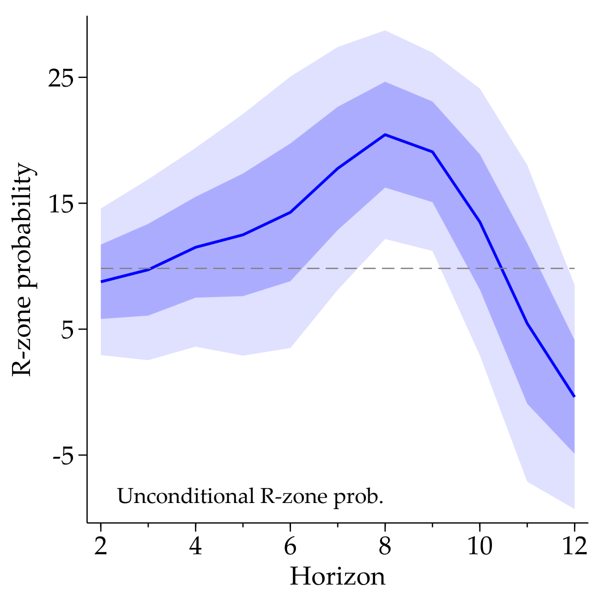
(b) *Full sample,*
$\overline{\text{stance}}$ lower by 100 bps



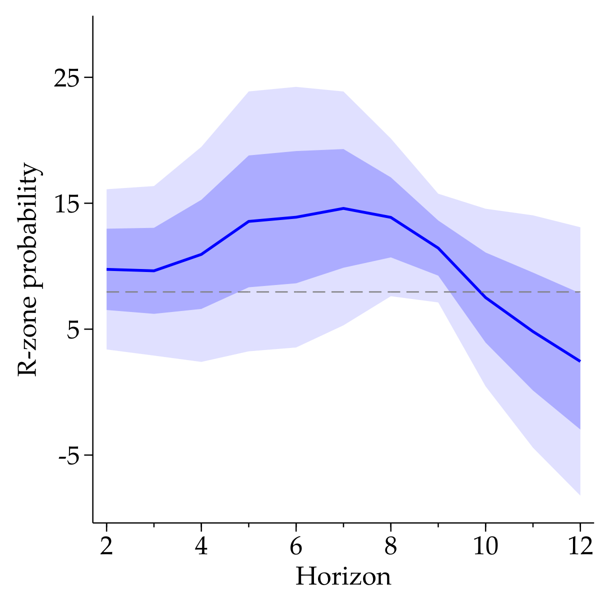
Notes: The same notes as in Figure 13 apply with one difference; we now consider housing finance R-zones, as in Figure 9.

Figure A13: *The effect of loose monetary policy on full-sample R-zones: second stage.*

(a) *Household sector,*
$\overline{\text{stance}}$ lower by 100 bps



(b) *Business sector,*
$\overline{\text{stance}}$ lower by 100 bps



Notes: The same notes as in Figure 13 apply with one difference; we now use all available data points of the JST Database.

A2 Adoption of the HLW approach to estimate the natural rate

An alternative, often-cited and often-used approach to estimate the natural rate of interest comes from Holston, Laubach, and Williams (2017; hereinafter HLW), which in turn builds on the seminal paper by Laubach and Williams (2003). We refer to HLW for a description of their three-stage model that builds on simplified versions of the two key equations of the New Keynesian framework. Contrary to the DGGT approach used in the main part of this paper, HLW estimate their model country by country. These countries are the United States, Canada, the Euro Area, and the United Kingdom. Their data covers the period from 1961Q1 (1972Q1 in the case of the Euro Area) to 2016Q3.

Our sample is characterized by a much broader coverage with respect to both time and space. A replication of the HLW approach to 18 countries for the period from 1870 to 2020 necessitates adjustments. First, we estimate the model in one stage rather than in three stages. That is, we do not start with reduced, auxiliary forms of the model which HLW use to derive Median Unbiased Estimates of λ_g and λ_z. Rather, we find maximum likelihood estimates of λ_g and λ_z by grid search. Second, we specify lower bounds for the grid and for the maximum likelihood estimation of $\sigma_{\tilde{y}}$ and σ_y in order to ensure a reasonable variation of trend growth g and other factors z. Third, we impose tighter constraints on the slopes of IS and Philips curves. The other adjustments concern the issues of outliers, world war years, and missing values. HLW were not confronted with these issues, so we have to make some additional decisions. As in the DGGT approach, we treat variables as missing whenever they exceed 30% in absolute values. Data for the world war years is ignored. We extrapolate over missing values in the Kalman filter and interpolate over missing values in the Kalman smoother as outlined in Durbin and Koopman (2001, p. 92–93).

Figure B1 reports our country-by-country two-sided estimates of the natural rate of interest r_{HLW} of this model. For comparison, it also shows r^* estimates of the DGGT approach.

As in the main text (see equation (4)), we define the stance of monetary policy as $\overline{stance}_{i,t} = \frac{1}{5}\sum_{k=0}^{4}\left(r_{t-k} - r^*_{HLW,t-k}\right)$. Our key result of this study is that a loose stance of monetary policy, i.e. a low value of $\overline{stance}$, increases the likelihood of the emergence of financial turmoil. We illustrated this finding in Figures 3 and 4. Figure B2 of this Appendix shows that this central result remains in place when using r^*_{HLW} to define the stance of monetary policy.

Figure B1: *By-country time series of r* for the two different approaches.*

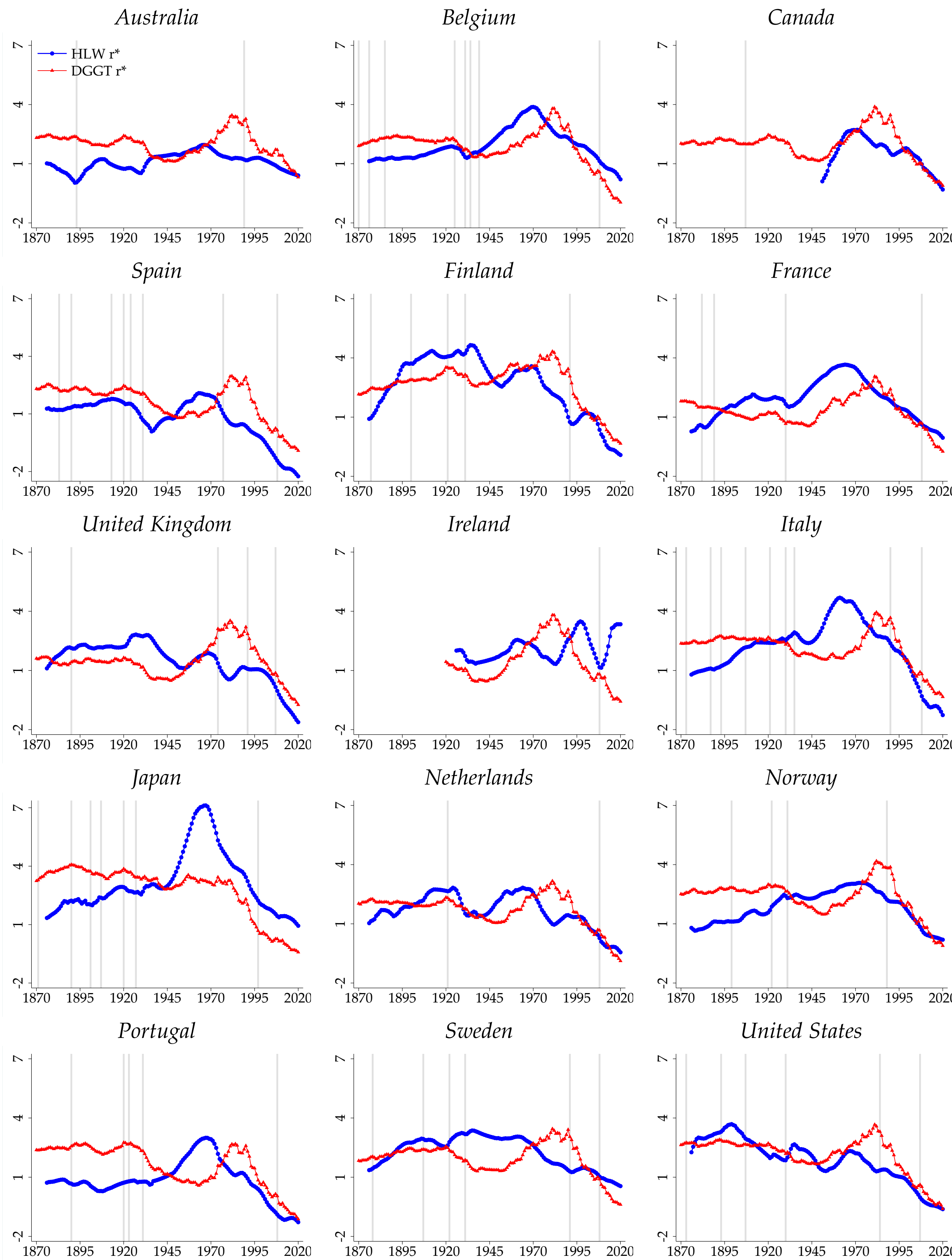
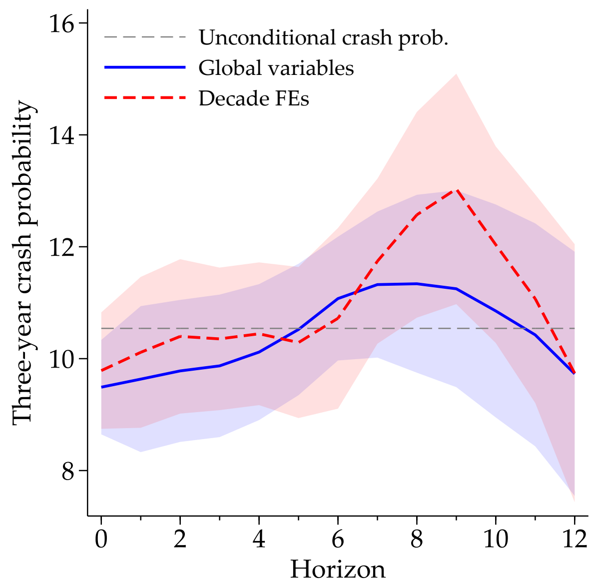
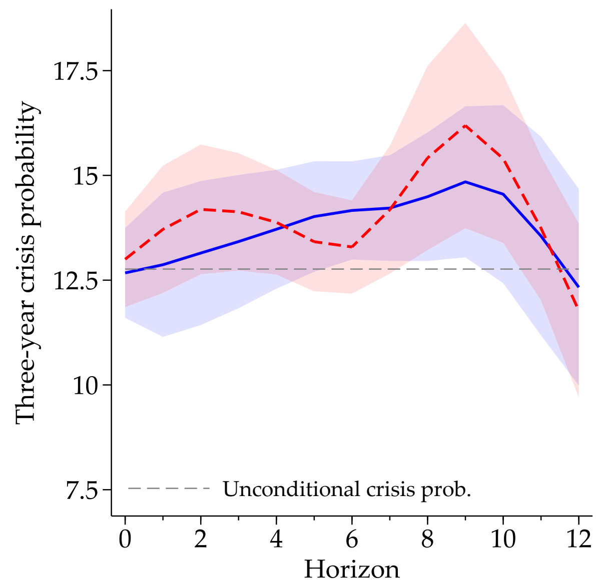


Figure B2: *The connection between loose monetary policy and financial turmoil: results with the Holston, Laubach, and Williams (2017) approach.*

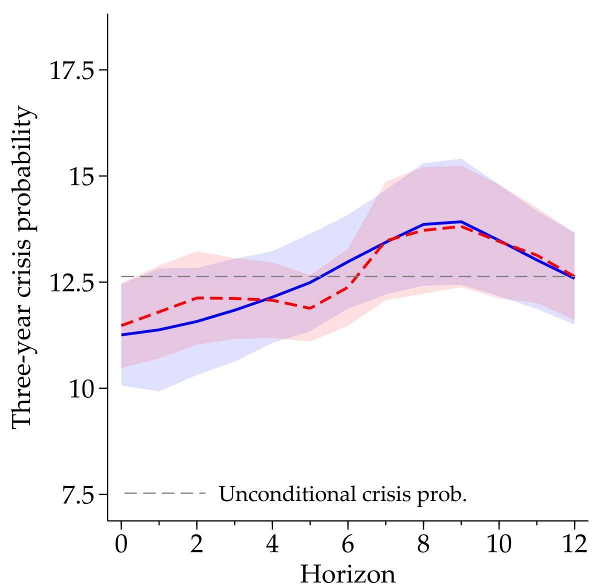
(a) *JST financial crises,*
$\overline{\text{stance}}$ lower by 100 bps



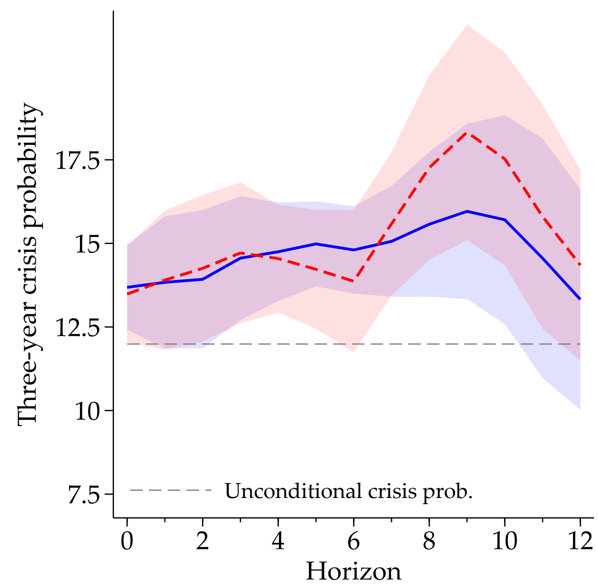
(b) Baron et al. (2021) *financial crises,*
$\overline{\text{stance}}$ lower by 100 bps



(c) Reinhart and Rogoff (2009) *financial crises,*
$\overline{\text{stance}}$ lower by 100 bps



(d) Baron et al. (2021) *bank equity crashes,*
$\overline{\text{stance}}$ lower by 100 bps



Notes: We re-estimate model (5) but define the natural rate and the stance of monetary policy based on the Holston, Laubach, and Williams (2017) approach. The solid blue lines show estimates based on the controls outlined in section 2.2. As an alternative specification, we replace the global control variables with decade fixed effects (dotted red lines). The figure shows estimates of $\{-100\beta^h\}_{h=0}^{12}$ relative to the unconditional full-sample three-year probability of the respective event and 95% confidence intervals based on Driscoll-Kraay (1998) standard errors with *ceiling*$(1.5 \times h)$ lags.