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Loose Monetary Policy and Financial Instability

MAXIMILIAN GRIMM
Federal Reserve Board
E-mail: maximilian.grimm@frb.gov

ÒSCAR JORDÀ
Federal Reserve Bank of San Francisco; University of California, Davis; and CEPR
E-mail: oscar.jorda@sf.frb.org

MORITZ SCHULARICK
Kiel Institute; Sciences Po Paris; and CEPR
E-mail: moritz.schularick@ifw-kiel.de

ALAN M. TAYLOR
Bank of England; Columbia University; NBER; and CEPR
E-mail: amt2314@columbia.edu

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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 in the medium term 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.

Key words: financial crises, crisis prediction, monetary policy, natural rate

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

1. INTRODUCTION

Does persistently loose monetary policy lead to financial fragility? And if so, why? Scholars and policymakers blamed loose monetary policy for the boom-bust cycle that culminated in the Global Financial Crisis (Geithner, 2009; Taylor, 2011). Moreover, they 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;

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Jordà et al., 2013) and political (Funke et al., 2016; Doerr et al., 2022) consequences of financial crises, there is up to now no systematic empirical study that analyzes the aggregate link between the stance of monetary policy and macro-level financial stability.

The answer to this question is not straightforward. There is an obvious tension: on the one hand lowering interest rates reduces the debt service burdens, stimulates the economy, and helps repair balance sheets, all of which point in the direction of lower rather than higher financial fragility. To get the opposite finding, significant adverse trends on the balance sheet must materialize, such as an endogenous increase in the quantity or decrease in the quality of the levered position—that is, more risk-taking. This paper seeks to understand which of these opposing forces dominates at the aggregate level. We are thus interested in the role that monetary policy potentially plays in triggering credit and asset price booms that are often seen as the precursors to financial instability.

On the micro-level, there are many theoretical and empirical studies on the connection between monetary policy and risk choices of financial institutions that we discuss below. Yet aggregate effects are not simply the sum of individual micro-level effects (Begenau and Stafford, 2023), so that the well-identified micro-level link between monetary policy and risk-taking does not necessarily allow us to draw conclusions about aggregate effects. On the macro side, a recent survey paper concluded that there is gap in knowledge about how micro-level risk-taking driven by monetary policy translates into measurable, aggregate-level, financial instability (Boyarchenko et al., 2022). On the theoretical side, we are only aware of recent work by Akinci et al. (2020) and Boissay et al. (2022) whose theoretical predictions match our empirical results, as we will show. Our paper therefore aims to fill this gap using macrofinancial data for the 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, among others, on the work of Del Negro et al. (2019).

As a first pass to explore the ideas in this paper, Figure 1 examines whether a negative *stance* precedes financial turmoil. This figure shows that in the years preceding the typical financial crisis, our stance measure is negative, suggesting that the stance of monetary policy might be too loose. For instance, panel (a) suggests that 4 years prior to the onset of financial crises, real rates are on average around 2 percentage points (pps) lower than the natural rate. Of course, other factors could be at play. Therefore, in the remainder of this paper, we go beyond the event-study exercise shown in Figure 1 and present formal evidence that when monetary policy is persistently loose, there are strong medium-term implications for macro-level financial stability.

An important attendant study by Jiménez et al. (2026) investigates the path of interest rates in the neighborhood of financial crises. Their approach shares elements with ours but differs in important respects. Like us, they start with the Jordà et al. (2017) Macrohistory Database and use the Jordà et al. (2020) trilemma-based instrument to investigate the dynamics of the *level* of interest rates leading up to financial crises. They complement their analysis using Spanish administrative data on banks from the

(a) Stance before Jordà et al. (2017) fin. crises (b) Stance before systemic financial instability

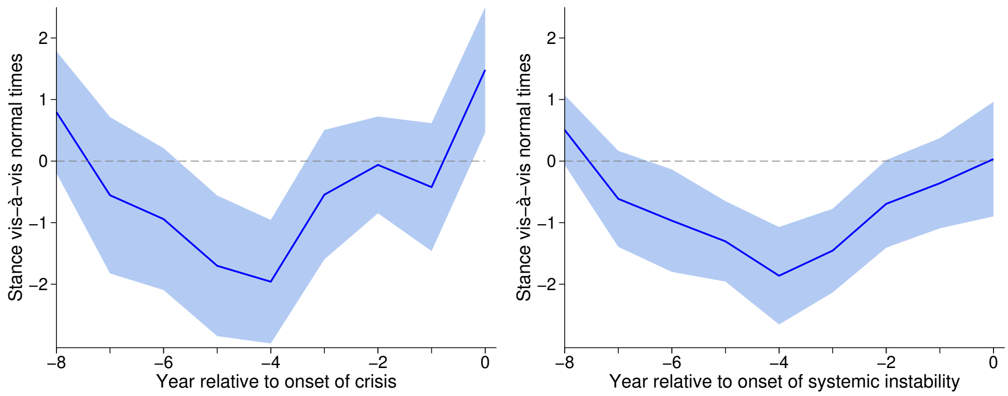

FIGURE 1

The stance of monetary policy before financial instability.

Notes: The figure shows estimates of $\{\beta^h\}_{h=0}^{8}$ of $stance_{i,t-h} = \beta^h y + \alpha_i^h + \alpha_{decade}^h + e_{i,t-h}$. The construction of our $stance_{i,t}$ measure in country i and year t is outlined in Section 2.1. y is specified in the panel titles. α refer to fixed effects. Shaded areas indicate 90% country-based cluster-robust confidence intervals. *Systemic financial instability*: union of the financial crisis and bank equity crash chronologies from Reinhart and Rogoff (2009), Jordà et al. (2017), and Baron et al. (2021).

Spanish credit register, available since 1995. Their key insight is that an asset-price boom concurrent with a period of low interest rates can collapse when interest rates suddenly jump, often leading to a financial crisis.

However, in their analysis, it is difficult to know whether the level of interest rates was consistent with the natural rate—and thus whether the main result is driven by the sudden jump in interest rates from any particular level, or whether this is true only when rates have deviated below their natural rate, regardless of what their absolute level might be. The distinction is not just semantic and matters greatly for central banks. In the former case, central banks have to worry about the risks of financial stability consequences whenever they raise rates. In the latter case, they only have to be worried about those risks if they have been pursuing a loose policy with rates below neutral for an extended period.

Hence, we provide the first systematic evidence that a prolonged accommodative stance—real rates kept below r^* for several years—significantly amplifies medium-term crisis risk, regardless of any particular pattern of cuts or hikes. Thus, while Jiménez et al. (2026) show that how central banks move nominal rates can influence financial stability, we establish that where policy is anchored—persistent real-rate deviations from neutrality—is a fundamental determinant of macro-financial vulnerability. Figure 1 illustrates how their findings align with ours: the positive slope of our stance measure in the immediate run-up to crises indicates that the late-cycle nominal-rate hikes they document also tighten the stance.

Our baseline OLS estimates suggest that a 1 pp lower average stance over a five-year window is associated with a 2.2 pps increase in the probability of a financial crisis occurring 5 to 7 years later, and a 3.2 pps increase 7 to 9 years ahead. Since the unconditional probability of experiencing a crisis in any 3-year window is 10.5 percent,

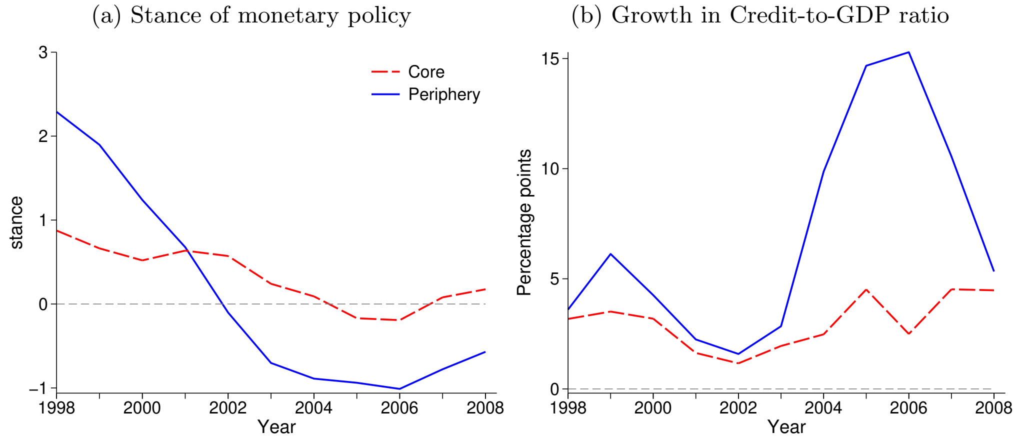

FIGURE 2

Stance of monetary policy in the pre-2009 eurozone.

Notes: Panel (a) shows the unweighted average of our stance measure, as defined in equation (2.4) of Section 2.1, for both the core and periphery countries of the eurozone. Panel (b) shows the corresponding unweighted averages of annual changes in the private-credit-to-GDP ratio. Core countries include Belgium, Denmark, France, Germany, and the Netherlands, while periphery countries consist of Ireland, Italy, Portugal, and Spain.

these effects are sizable. Moreover, these results are robust to alternative measures of the policy stance as well as to alternative definitions of financial instability.

We then explore *why* excessively loose monetary policy triggers financial instability. In particular, we look into the probable channels through which an accommodative stance translates into increased financial fragility, focusing on the risks of overheating in credit and asset markets. Here, using the "Red-zone" definitions of Greenwood et al. (2022), we find that loose-for-long monetary policy is associated with a buildup in asset prices and credit growth. In line with the existing macro-finance literature, this then helps predict financial crises.

In the second half of the paper, we introduce an instrumental variable strategy aimed at validating our reduced-form findings. Identification requires exogenous variation in the triggers of credit and asset price booms and financial crisis risk, as has been recognized in the literature. Our novel instrumental variable approach is based on the trilemma of international finance (Obstfeld and Taylor, 2004) and has been used to study the effects of monetary policy by, e.g., Jordà et al. (2020). As in Jordà et al. (2020), the differences between the OLS and IV results are stark, lending support to a causal reading conditional on the instrument's exclusion restriction.

An illustration of how trilemma-based identification works, and how loose monetary policy causes booms in financial markets that ultimately bust, is provided by the now-familiar tale of the eurozone in the 2000s. During this period, the countries within the euro area experienced heterogeneous growth rates. While the periphery was booming, the core exhibited only moderate growth. If monetary policy reacted to economic conditions in the core only, then the European Central Bank's interest rate decisions were too accommodative for periphery countries like Spain and Ireland. Over time, this should imply a looser stance of monetary policy in the periphery compared to the core. Indeed, the literature argues that in the years before the Global Financial Crisis, the European

Central Bank's monetary policy stance was in line with macroeconomic conditions in the core countries but too accommodative for the periphery (Nechio, 2011; Lothian, 2014). Our stance measure matches up with these judgments from independent analyses, as shown in Figure 2 (a), providing additional confidence from a specific historical episode. The panel plots the average of our stance measure for the core and periphery countries of the eurozone. For the eurozone's periphery countries, our stance estimate features a sharp downward trend and is *too loose* over the years preceding the crisis. In contrast, the fall of the stance in the core was only moderate. For these core countries, our estimates do not indicate a significant deviation from a neutral stance of monetary policy. This divergence in policy stance is what our instrument captures, as outlined below. Panel (b) of Figure 2 illustrates that precisely those countries that were confronted with too loose a stance of monetary policy experienced a rapid credit boom in the mid-2000s. This exemplifies the link between a loose policy stance and credit market overheating, a relationship this study will systematically establish across time and space.

How much should one trust our results? Like any empirical study based on a single instrument for a single endogenous variable, the validity of the instrument is statistically unverifiable. One has to rely on economic arguments and here we think we stand on solid footing. The stark differences between the OLS and IV estimates further support this conclusion.

Nevertheless, one can also suspect that the exclusion restriction is violated. While this is also statistically unverifiable, there are reasonable economic arguments challenging this assumption. Hence, we explore the robustness of our results using the methods in Conley et al. (2012). Once again, our findings seem robust to economically plausible violations of the exclusion restriction. Thus, while it is important to be modest about any empirical analysis, we are prudently confident in the results that we report.

In the last part of this paper, we show that prolonged accommodative 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 et al. (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 et al., 2019, 2022). This suggests an intertemporal trade-off between real activity today and higher instability risk down the road (Drehmann et al., 2012). 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 (2024).

Theory and evidence at the micro level: risk-taking and reach for yield. The stance of monetary policy likely affects risk taking (Bauer et al., 2023). As already noted earlier, several empirical studies show that when monetary policy is loose over an extended period, financial institutions become less profitable (Claessens et al., 2018) and engage in riskier investment decisions (Maddaloni and Peydró, 2011; Altunbas et al., 2014; Chodorow-Reich, 2014; Jiménez et al., 2014; Ioannidou et al., 2015; Hau and Lai, 2016; Dell'Ariccia et al., 2017; Di Maggio and Kacperczyk, 2017; Paligorova and Santos, 2017; Choi and Kronlund, 2018). To the best of our knowledge, only Altunbas et al. (2014) attempt 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 et al. (2019) from which to calculate the stance of monetary policy.

When financial institutions such as insurance companies, who often face fixed long-term commitments, are confronted with "extremely accommodative monetary policy" (Rajan, 2005, p. 517), they have incentives—or are even required—to "search for yield" and hence risk (Rajan, 2005). More recently, Campbell and Sigalov (2022) have rationalized search-for-yield behavior arising from a sustainable spending constraint. Hanson and Stein (2015) argue that yield-oriented investors—who place particular emphasis on current returns—increase their portfolio's maturity risk when short-term rates fall, and empirically support this argument; after expansionary monetary policy shocks, the maturity of banks' security holdings rises. Other studies have established a theoretical link between low rates and investors' leverage and risk exposure by focusing on yield-oriented investors (Hanson and Stein, 2015), the cost of liquidity buffers (Drechsler et al., 2018), the benefit of capital as a signaling device (Dell'Ariccia et al., 2014), and monitoring incentives (Dell'Ariccia et al., 2014; Martinez-Miera and Repullo, 2017; Heider and Leonello, 2021). If a loose stance of monetary policy creates abundant liquidity, it may 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, p. 350). 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 et al. (2019) find that individuals starting their 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.

Mechanisms at the macro-level: aggregate consequences of a loose stance. Credit and asset price booms have been identified by the literature as harbingers of financial turmoil. Elevated credit growth (Mendoza and Terrones, 2008; Schularick and Taylor, 2012; Jordà et al., 2016), house price booms (Jordà et al., 2015a), and the interaction of credit and asset price booms (Borio and Lowe, 2002; Jordà et al., 2015b; Greenwood et al., 2022) are important short-run predictors of financial crises. Similarly, credit expansions predict bank equity crashes (Baron and Xiong, 2017).

However, little is known about the triggers of booms in financial markets (Mian and Sufi, 2018, pp. 50–52).[1] Moreover, the literature examining the (medium-term) relationship between the stance of monetary policy and systemic financial instability is virtually nonexistent, with only a few notable exceptions on the empirical (Jiménez et al., 2026) and theoretical (Akinci et al., 2020; Boissay et al., 2022) side. Kashyap and Stein (2023) stress that "there is limited evidence that it is specifically monetary-policy induced changes in credit growth and risk premiums——as opposed to changes driven by

1. Jordà et al. (2015a) provide evidence that a loosening of monetary conditions leads to an increase in credit and asset prices. Bianchi et al. (2022) show in the context of the U.S. that long-lasting shifts in asset valuations coincide with shifts in the Fed's conduct of monetary policy.

other factors—–that create this economic vulnerability" since "establishing such a link is challenging, and more research on this specific issue would be valuable." (p. 63) The contribution of our study lies in addressing this gap.

Our empirical analysis builds on Boissay et al. (2022), who analyze the macro-financial consequences of lower-for-longer monetary policy from a theoretical perspective. In their model, financial crises occur when central banks keep policy rates too low for too long. This persistently loose monetary policy, caused by a series of expansionary monetary policy shocks, triggers an investment boom that eventually culminates in a collapse of credit markets. Our empirical framework is motivated by Boissay et al. (2022); we explicitly examine the consequences of a *persistently* loose stance of monetary policy caused by expansionary shocks. Our results support their theoretical findings; protracted loose monetary policy fuels a boom in financial markets and leads to large-scale financial disruptions. Similarly, in the model of Akinci et al. (2020), persistent reductions in real interest rates lead to improved financial conditions and higher risk-taking in the short term, which triggers higher leverage in the medium term that ultimately reduces policy space and heightens financial fragility further down the road.

This study also contributes to the literature documenting the long duration of financial cycles (Drehmann et al., 2012; Claessens et al., 2012; Borio, 2014), which "depend on the policy regimes in place" (Borio, 2014, p. 185) and tend to peak around the onset of financial crises (Drehmann et al., 2012; Borio, 2014). We identify a significant medium-term relationship between the stance of monetary policy and crisis risk, indicating that monetary policy is an important driver of the financial cycle.

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. For instance, the stance of U.S. monetary policy in the 1970s was *loose* (Clarida et al., 2000), but 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 personal views 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 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 (Borio and Lowe, 2002; Boissay et al., 2022). A case in point is the Global Financial Crisis, which was also preceded by accommodative monetary policy, but not by high inflation. 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, as illustrated above. Eventually, maintaining too loose a stance generates a set of possibly perverse incentives, discussed above, regardless of its cause.

2. FRAMEWORK

Our empirical analysis is based on the latest 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 at annual frequency. The database is described in Jordà et al. (2017).[2] 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 about 2,500 country-year observations.

Of course, not all variation in real policy interest rates is due to monetary policy—some of the dynamics of interest rates may be due to secular trends in the natural rate of interest arising for other fundamental reasons, such as demographics, productivity growth, and so on (see, e.g., Rachel and Summers, 2019). Woodford (2003, p. 248) defines the natural rate 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).

Thus, 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, $r \equiv r^* + stance$. 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. Hence, though we use the wording "loose monetary policy" throughout this study, stance does not have to be only a direct function of intentional domestic monetary policy choices (rate shocks) by the central bank, it could also reflect their mismeasurement of a changing r^*, for example.

2.1. *Estimating the natural rate and stance*

Rather than developing a new method, we follow the literature which has proposed several established methods to estimate the natural rate. We estimate *stance* by extending the work of Del Negro et al. (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.

Several methods have been proposed in recent years to identify r^*. The resulting estimates of r^* vary to some extent. Except for the DGGT approach, all these methods possess at least one of the following three characteristics: a focus on the United States, samples drawn from recent decades, and estimates based on data at higher-than-annual frequency. For these reasons, it is natural to extend DGGT as our baseline method. As

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

a robustness check, we also replicated the Holston et al. (2017) approach and extended it to all 18 countries of our study. In Appendix B, 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 approach to measuring the natural rate.

Furthermore, unlike many studies that estimate r^* (e.g., Lubik and Matthes, 2015; 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. 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 obtained using Bayesian estimation methods.

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 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

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

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

As in DGGT, we assume that the state variables follow a VAR process with one lag, noting that we use annual data. When allowing for two lags, estimates of the long-run trends in interest rates, as well as our main findings, are similar, as demonstrated below. 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). The structure of $\boldsymbol{\Lambda}$ in (ME), on the other hand, is restricted. DGGT's asset pricing model begins with an international investor's pricing equations for bonds of various countries, from which long-run no-arbitrage conditions on countries' interest rates are derived. In the empirical model, these theoretically derived conditions imply that the trends of country i's short-term real rate and term spread are equal to $\bar{r}^w_t + \bar{r}^i_t$ and $\bar{ts}^w_t + \bar{ts}^i_t$, respectively. However, since there is no equivalent to a marginal international investor that ensures the absence of long-run arbitrage opportunities in international bond markets, we cannot infer conditions from the asset pricing model that discipline the behavior of inflation. Therefore, country-specific loadings λ^π_i on the world inflation trend are allowed. That is, trends in inflation are not the sum of global and country-specific trends but are equal to $\lambda_i^\pi \bar{\pi}_t^w + \bar{\pi}_t^i$.[3]

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

$$R_{i,t} = \bar{r}_t^w + \bar{r}_t^i + \lambda_i^\pi \bar{\pi}_t^w + \bar{\pi}_t^i + \tilde{R}_{i,t}, \tag{2.1}$$

$$R_{i,t}^L = \bar{r}_t^w + \bar{r}_t^i + \bar{ts}_t^w + \bar{ts}_t^i + \lambda_i^\pi \bar{\pi}_t^w + \bar{\pi}_t^i + \tilde{R}_{i,t}^L, \tag{2.2}$$

$$\pi_{i,t} = \lambda_i^\pi \bar{\pi}_t^w + \bar{\pi}_t^i + \tilde{\pi}_{i,t}. \tag{2.3}$$

Then, following DGGT, we interpret $r_{i,t}^* \equiv \bar{r}_t^w + \bar{r}_t^i$ as the natural rate of interest of country i.

After subtracting equation (2.3) from equation (2.1), 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. Therefore, we 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.[4] 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,

$$\overline{stance}_{i,t} = \frac{1}{5} \sum_{k=0}^{4} \left(r_{i,t-k} - r_{i,t-k}^* \right). \tag{2.4}$$

This stance measure is our main object of interest. It represents the relevant independent variable in all econometric models we estimate. Robustness checks reported later show

3. Notice that posterior draws of λ_i^π are closely centered around 1 for most countries as shown in Appendix Figure A2. We also estimated two alternative specifications of the state-space model that impose different restrictions on $\mathbf{\Lambda}$. In the first robustness check, we outright imposed $\lambda_i^\pi = 1$. In the second sensitivity check, we estimated *both* λ_i^π and country-specific loadings λ_i^r on the world short-term real rate trend. In the second sensitivity check, posterior draws of λ_i^r are closely centered around 1 for all countries (Appendix Figure A3), suggesting that the long-run no-arbitrage conditions on interest rates are supported by the data. These two alternative specifications of the state-space model do not significantly affect either the estimates of the latent variables or our main findings as verified in Appendix Figures A4 and A5, respectively.

4. 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 et al. (2022) introduce a model that features such long-lasting deviations from monetary neutrality after regime changes in the conduct of monetary policy.

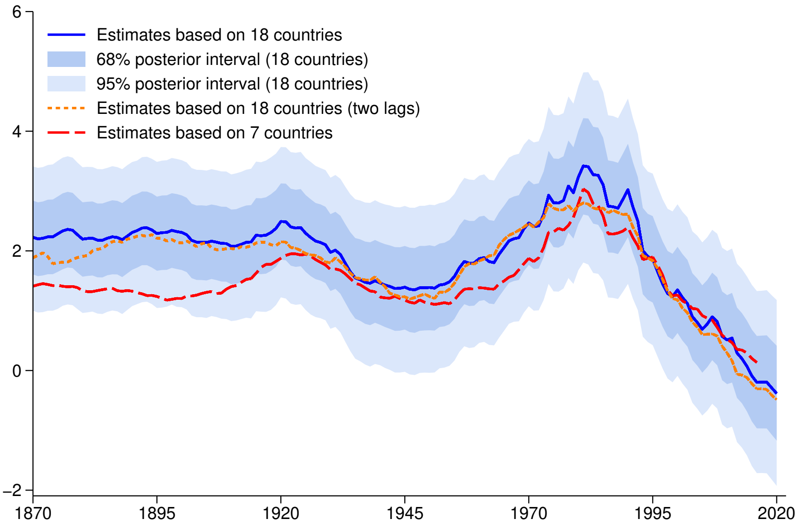

FIGURE 3

Trends in the global real rate.

Notes: The solid blue and dashed red lines represent posterior medians of $\bar{r}_t^w$ for our model with 18 countries (which uses data until 2020) and for the original DGGT model with 7 countries (which uses data until 2016), respectively. The dotted orange line shows posterior medians of $\bar{r}_t^w$ when allowing for two lags in the VAR in (SE), keeping all other specifications unchanged.

that the exact number of years over which we take the average does not affect our results significantly. Additionally, we will validate our findings by substituting the ex-post real interest rate in equation (2.4) with two alternative measures of ex-ante real rates.

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 $\bar{r}_t^w$ 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.[5] 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.

The solid blue and dashed red lines in Figure 3 show estimated world trends of the natural rate of interest, that is posterior medians of $\bar{r}_t^w$, for both our model with 18 countries and the original DGGT model with 7 countries. The two estimated series

5. 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 et al. (2022) who replicate the DGGT approach for a sample of 31 countries.

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" (p. 249) implying smaller country-specific trend components. This finding carries over to our 18-country model; before the late 1970s, country-specific trends were larger.[6] 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.[7] We can note in this context that Cesa-Bianchi et al. (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 et al., 2017; Jordà and Taylor, 2019). According to our definition of stance in equation (2.4), many countries were therefore not in an environment characterized by exceptionally loose monetary policy in the 2010s.

The estimates of $\bar{r}_t^w$ look similar when allowing for two lags in the VAR in (SE). This similarity is illustrated by the dotted orange line in Figure 3. Its dynamics over time resemble the estimates of our baseline model and the original DGGT model, both of which only allow for a one-year lag in the VAR. Below, we confirm that our main empirical results also remain in place when defining our stance measure based on the estimation of a VAR with two lags.

Full country-by-country estimates are shown in the Appendix. Figure A1 shows estimated 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 (2.4) for all 18 countries of our sample. Whenever applicable, we also show replicated country trends from the 7-country DGGT model. Gray vertical bars indicate the onset year of financial crises.

To sum up, extracting the trend matters. Today's investors live in a low real rate world, but their predecessors did not. For example, the red line in the bottom-right panel of Appendix 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 5.77% in the 1870s due to high natural rates (blue line). In contrast, while $r_{USA} - r^*_{USA}$ was on average negative in the 2010s (-1.03%), 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

6. 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 0.49. Also see Appendix Figure A1.

7. In the post-1977 period, the cross-country average of real interest rates is 1.69% for the G7 and 1.68% for the other 11 countries. Corresponding averages for the period up to 1977 are 0.24% and 2.12%.

be written as a Jordà (2005) local projection, or LP,

$$B_{i,t+h} = \beta^h \overline{stance}_{i,t} + \alpha_i^h + \mathbf{\Gamma}^h \mathbf{X}_{i,t} + \sum_{k=1}^{5} \delta_k^h b_{i,t-k} + u_{i,t+h}, \qquad h = 0,\ldots,H. \tag{2.5}$$

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). Our independent variable of interest, $\overline{stance}$, may have an endogenous component. We address this concern in two ways. We start by estimating model (2.5) by OLS subject to a comprehensive set of control variables that could potentially confound our estimates. Subsequently, we validate our results within an IV approach.

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 (2.5), 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, 2025; 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, the stance of monetary policy affects the slope of the yield curve, which in turn is a strong short-term predictor of financial crises (Parker and Schularick, 2021; Bluwstein et al., 2023).

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.

Ultimately, the β^h are 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. Statistical inference allows for complex structures of the error term by computing Driscoll-Kraay (1998) standard errors that are robust to heteroskedasticity as well as 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 above to construct 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 $\overline{stance}$, are subject to first-stage estimation uncertainty, as seen in Figure 3. 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.[8]

A second potential issue is that the first-stage uncertainty is akin to a measurement error that biases the second-stage coefficient estimates towards zero. In Section 4, we use 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 accommodative monetary policy and the likelihood of financial turmoil down the road.

3. LOOSE MONETARY POLICY AND FINANCIAL INSTABILITY

3.1. *Financial crises*

This section asks whether periods of persistently loose monetary policy are more crisis-prone. Hence, 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 (2.5) with: $B_{i,t} = \max\{crisis_{i,t}, crisis_{i,t+1}, crisis_{i,t+2}\}$;[9] 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 et al., 2021). We will ensure that results hold for alternative 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 $B_{i,t} = b_{i,t} = crisis_{i,t}$.

8. We show bootstrap confidence intervals for our main results on the relation between the *stance* and crisis risk (Figures 4 and 11). We also verified the robustness of our other results to this first-stage estimation uncertainty.

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

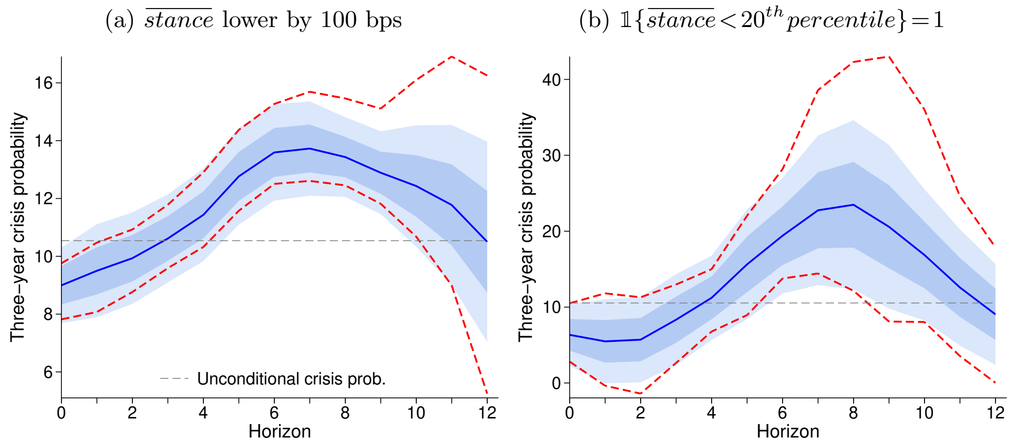

FIGURE 4

The connection between loose monetary policy and financial crises.

Notes: Panel (a) shows estimates of $\{-100\beta^h\}_{h=0}^{12}$ of equation (2.5) 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 $\overline{stance}_{i,t}$ with 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. Shaded areas indicate 95% (light) and 68% (dark) confidence intervals based on Driscoll-Kraay (1998) standard errors with $ceiling(1.5 \times h)$ lags. The red dashed lines denote 95% confidence intervals based on the bootstrap procedure described in Section 2.3.

Figure 4 presents our main result. Panel (a) shows the estimated change in crisis risk, relative to the full-sample unconditional three-year crisis probability (gray dashed line), when $\overline{stance}$ as defined in equation (2.4) is lower by one percentage point. Panel (b) displays the change in crisis risk when $\overline{stance}$ is in the lowest quintile. Shaded areas indicate confidence intervals based on analytic standard errors. The red dashed lines represent bootstrap confidence intervals that account for first-stage uncertainty as discussed in Section 2.3.

Consider panel (a) first. Loose monetary policy predicts elevated medium-term financial crisis risk 5 to 10 years ahead.[10] This empirical finding is not just statistically significant (and robust, as we will show), it is also economically relevant. When rates are 1 pp lower than the natural rate, on average over five years, then financial crisis risk increases by 2.2 pps 5 to 7 years ahead and by 3.2 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 (indicated by the gray dashed line in Figure 4) these estimates are quantitatively important. And we will find even larger effects below using our preferred IV specification.

Notice that the sample mean of $\overline{stance}$ is approximately zero. 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

10. 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 (Claessens et al., 2012; Drehmann et al., 2012).

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

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

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

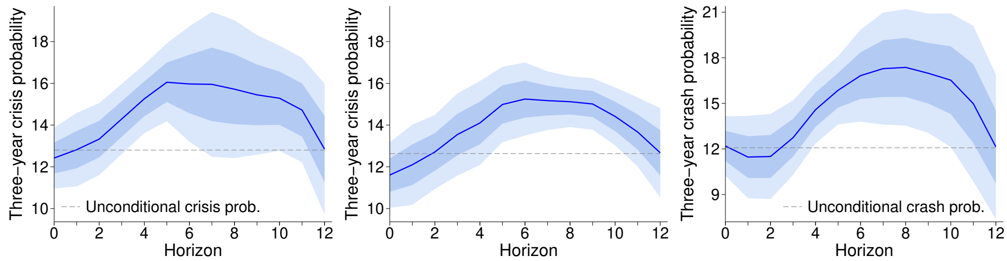

FIGURE 5

The connection between loose monetary policy and alternative financial instability event indicators.

Notes: The same notes as in Figure 4 apply.

to about 17.9% when $\overline{stance} = -2.4\%$ (which is the 20^{th} percentile of the pooled country-year $\overline{stance}$ distribution) 6 to 8 years ahead.

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 with a binary one in model (2.5). 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 4. Impulse responses exhibit a similar pattern to those in panel (a).

Robustness: Symmetry. One may ask whether the effects are asymmetric. Specifically, does tight monetary policy have the opposite effects of loose monetary policy? We do not find evidence of such asymmetry. To check this, we divide the pooled country-year $\overline{stance}$ distribution into five quintiles ($Q^1_{i,t} = \mathbb{1}\{\overline{stance}_{i,t} < 20^{th} percentile\}$, $Q^2_{i,t} = \mathbb{1}\{20^{th} percentile \leq \overline{stance}_{i,t} < 40^{th} percentile\}$, ...) and then estimate the modified version of (2.5),

$$B_{i,t+h} = \beta^h_1 Q^1_{i,t} + \beta^h_2 Q^2_{i,t} + \beta^h_4 Q^4_{i,t} + \beta^h_5 Q^5_{i,t} + \alpha^h_i + \mathbf{\Gamma}^h \mathbf{X}_{i,t} + \sum_{k=1}^{5} \delta^h_k b_{i,t-k} + u_{i,t+h}, \quad (3.6)$$

for $h = 0,...,H$. Detailed results are presented in Appendix Figure A6, and show that while a stance within the lowest quintile of its pooled country-year distribution significantly amplifies medium-term crisis risk ($\beta_1 > 0$), a stance in the highest quintile reduces it ($\beta_5 < 0$) by a similar magnitude.

Robustness: Alternative chronologies. 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 5. 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 et al. (2021).[11] In panel (c), we replace the crisis indicator with the bank equity crash indicator from Baron et al. (2021). This binary indicator is on in years when bank equity returns are −30% or less. 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 gray dashed line.

Robustness: Crisis window. 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 was 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 (2.5) for different values of F. Results of this exercise are shown in Appendix Figure A7. 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 5% level. Similarly, our main result shown in Figure 4 also holds when averaging $r - r^*$ over windows of different lengths. This is illustrated in Appendix Figure A8.

Robustness: Lag length. As explained in Section 2.1, we follow DGGT by constraining the lag length of the VAR process in (SE) to one year. To ensure that our results do not depend on this restriction on (SE), we redefine $\overline{stance}$ using r^* estimates obtained from a VAR model with two lags. All other specifications of the DGGT model remain unaltered. Estimation results of model (2.5) based on this redefined stance measure are similar, as confirmed in Appendix Figure A9 (a).

Robustness: Logit. Throughout this study, our findings are derived from the estimation of a simple linear probability model which makes the interpretation of coefficients straightforward. However, it is more natural to model the probability of a binary dependent variable with a logit model, for example. This does not have much impact on our results, as shown in Appendix Figure A10.

Robustness: Business cycle. One possible worry is that a low stance simply reflects periods of economic expansion. 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. Two features allay this concern. 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. Thus, our results show that a loose stance predicts financial crises once the state of the business cycle is accounted for. Second, we can show that $\overline{stance}$ is not a signal for normal business-cycle downturns; it only predicts recessions associated with financial turmoil. In Appendix Figure A11 we re-estimate our main econometric model (2.5) for a different dependent variable: a binary indicator with one of two mutually exclusive and exhaustive types of recessions defined

11. 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).

as in Jordà et al. (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 recessions that go hand-in-hand with a crisis.

Robustness: Time fixed effects. 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 Appendix Figure A12, respectively.

Robustness: Alternative $\boldsymbol{r}^*$*.* Furthermore, the relationship between the stance of monetary policy and financial instability risk remains alive and well for the post-WWII period (Appendix Figure A13) and when using the Holston et al. (2017) approach to estimate the natural rate of interest (Appendix 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 B, which in turn are reflected in the heightened estimation uncertainty apparent in Appendix Figure B2. Nevertheless, point estimates exhibit a similar shape as in Figures 4 (a) and 5.

Robustness: Ex-ante stance. Finally, we assess whether our findings are sensitive to constructing $\overline{stance}$ based on ex-post real rates. Historical datasets, including ours, do not contain inflation expectations, as they were not available in the distant past either from surveys or market-based measures. However, OECD inflation forecasts for various countries for several decades have been recently digitized by Adam et al. (2023). Their inflation forecast dataset covers all 18 countries of the JST Database and starts, for some countries, as early as 1965. In total, the OECD's inflation forecasts cover 885 observations of the JST Database.

In Appendix Figure A14, the solid blue and dotted red lines show realized inflation and OECD inflation forecasts, respectively, for the United States over the past six decades. As a second established approach, different from using professional forecasters' expectations, we proxy inflation expectations with constant-gain learning and construct an alternative measure of inflation expectations as a discounted moving average (DMA),

$$\tau_t^{CPI} = (1-\nu)\sum_{i=0}^{t-1}\nu^i \pi_{t-i}. \tag{3.7}$$

We set $v = 0.987^{12}$ and truncate the sum in equation (3.7) at $N = 120/12$ years. This specification and notation, as well as the choices of v and N, follow Cieslak and Povala (2015, p. 2865). The resulting series for the United States, represented by the dashed orange line in Appendix Figure A14, mirrors the one presented in Cieslak and Povala (2015), which in turn “comoves closely with inflation expectations (perceived inflation target) underlying the Federal Reserve Board's FRB/US model” (Cieslak and Povala, 2015, p. 2867). Denote by $\pi^e_{t,t+1}$ inflation expectations formed in year t for year $t+1$. Equipped with two measures of expected inflation, we define the stance of monetary policy as

$$\overline{stance}_{i,t} = \frac{1}{5}\sum_{k=0}^{4}\left(R_{i,t-k} - \pi^e_{i,t-k,t-k+1} - r^*_{i,t-k}\right). \tag{3.8}$$

As before, R refers to the short-term nominal interest rate. Notice that equation (3.8) deviates from equation (2.4) solely in the specification of inflation. Here, we use one of the two aforementioned measures for expected inflation, whereas in all other parts of this study, ex-post inflation is used to determine the stance of monetary policy. We proceed to re-estimate our baseline model, summarized in Section 2.2. All specifications, including the selection of control variables and the construction of the instrumental variable, remain unchanged. Results appear in panel (a) of Appendix Figure A15 and show that our key finding is robust to using alternative measures of expected inflation in place of ex-post inflation.

3.2. *The level of interest rates versus the stance*

Another concern is perhaps the idea that it might be the *level* of the interest rate, as opposed to, or in addition to, the stance that matters for crisis prediction. A priori, this is unlikely.

Recall that r^* by definition equates saving and investment in an economy without nominal rigidities. That is, in *relative terms*, 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 et al. (2022).

The natural rate r^* is the long-run trend of the real rate. We can therefore also think of r^* as a proxy for the real rate on long-term fixed-interest liabilities of financial institutions. But when observed real rates are lower than r^*, financial institutions with a large share of such long-term liabilities may feel 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 et al., 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 et al., 2022). Our framework allows us to do exactly that.

To do this, Figure 6 shows estimates of model (2.5) 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. This evidence suggests that it is the conduct of monetary policy, rather than the interest rate level itself, that plays the dominant role in our main results.

3.3. *Credit and asset market overheating*

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

(a) $\frac{1}{5}\sum_{k=0}^{4} r_{i,t-k}$ lower by 100 bps (b) $\mathbb{1}\{\frac{1}{5}\sum_{k=0}^{4} r_{i,t-k} < 20^{th} percentile\} = 1$

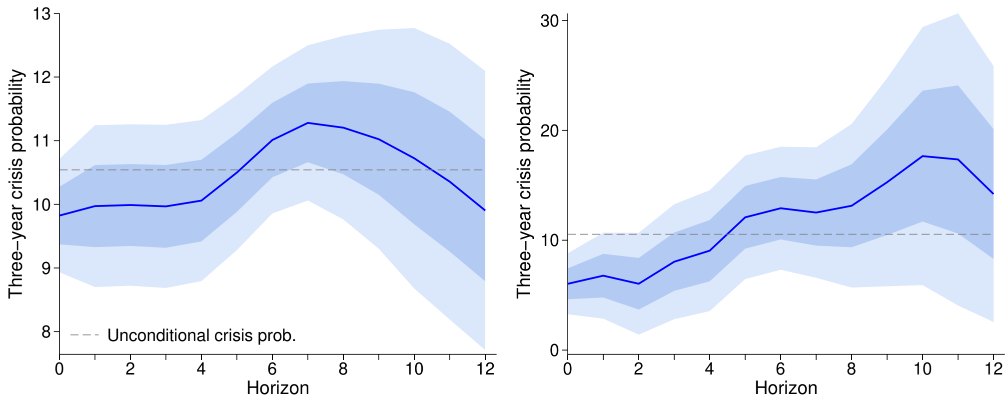

FIGURE 6

The connection between real interest rates and financial crises.

Notes: The same notes as in Figure 4 apply with one difference; the only interest variable included in model (2.5) is now $\frac{1}{5}\sum_{k=0}^{4} r_{i,t-k}$.

harbingers of financial crises (Jordà et al., 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 et al. (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.

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 (2.5), keeping the vector of control variables $\mathbf{X}$ unchanged, but redefining the binary dependent variable B to refer to R-zones, $B_{i,t} = R\text{-}zone_{i,t}$, where R-zones are defined as in Greenwood et al. (2022), specifically

$$\begin{aligned} &High\text{-}Debt\text{-}Growth_{i,t} = \mathbb{1}\{\Delta_3 (Debt/GDP)_{i,t} > 80^{th} percentile\}, \\ &High\text{-}Price\text{-}Growth_{i,t} = \mathbb{1}\{\Delta_3 \left(\log Price_{i,t}\right) > 66.7^{th} percentile\}, \\ &R\text{-}zone_{i,t} = High\text{-}Debt\text{-}Growth_{i,t} \times High\text{-}Price\text{-}Growth_{i,t}. \end{aligned} \tag{3.9}$$

The High-Debt-Growth indicator, $High\text{-}Debt\text{-}Growth_{i,t}$, takes on a positive value if the change in country i's debt-to-GDP ratio from year $t-3$ to year t is in the top quintile of its distribution. The distribution is derived from the full country-year panel that we are using. The High-Price-Growth indicator is defined in a similar way for the top tercile.[12] By definition, the economy is then in an R-zone when both debt and prices rise relatively strongly from a historical point of view.

12. We confirmed the robustness of our main findings, presented below, by substituting three-year changes (i.e., Δ_3) with one-year changes (i.e., Δ_1) in equation (3.9).

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	12.82	12.58
Stock prices	22.82	22.22

As in Greenwood et al. (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. As a reference, Greenwood et al. (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 et al., 2017; Müller and Verner, 2024).

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 additional 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 is turned on, strong 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 et al. (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 Figure 4 (a) indicates that the connection between $\overline{stance}$ and crisis risk is strongest at a horizon of 7 years. Similarly, Figure 7 (a) indicates that the association between $\overline{stance}$ and R-zone risk is most pronounced after 7 years. Consequently, our combined findings suggest that a loose stance of monetary policy in year t is associated with both rapid credit and asset price growth between $t+4$ and $t+7$, as well as an increased crisis risk between $t+7$ and $t+9$. While the short-term association

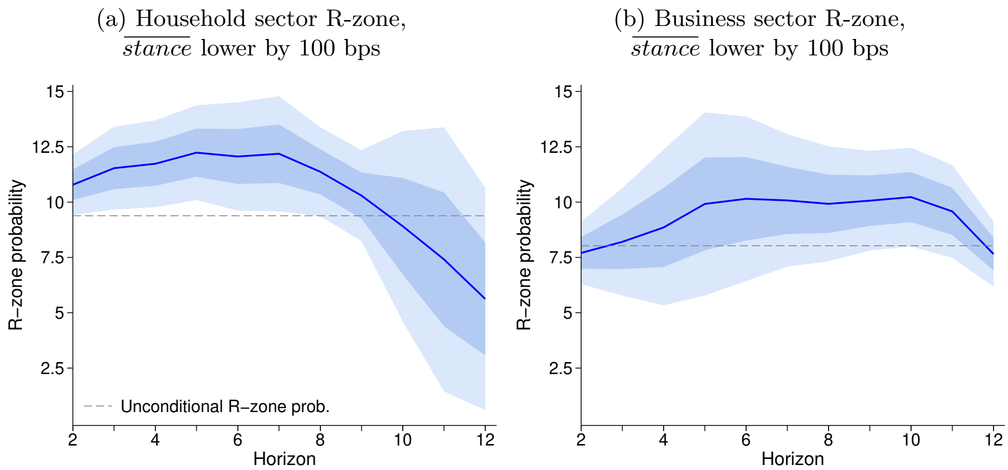

FIGURE 7

The connection between loose monetary policy and post-WWII R-zones.

Notes: The figure shows estimates of $\{-100\beta^h\}_{h=2}^{12}$ of model (2.5) with R-zones as defined in equation (3.9) 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.

between a loose stance and crisis risk is *negative*, the likelihood of experiencing rapid credit and asset price growth rises immediately and remains elevated after excessively loose monetary policy. As a result, the cumulative probability that the economy has entered an R-zone between years t and $t+7$ is considerably higher than suggested by the individual point estimates.

Figure 7 suggests that R-zone risk stays high for several years even though the risk of financial crises, which typically go hand-in-hand with collapsing financial markets, also rises. How can this be? Medium-term crisis risk is not deterministic—it is just more likely after persistently loose monetary policy. Overheated financial markets push the economy closer to the financial stress region, exposing it to possibly greater dangers from negative technology shocks and contractionary monetary policy shocks (Akinci et al., 2020; Boissay et al., 2022; Jiménez et al., 2026). Our results are consistent with a view that an abrupt reversal of rates can be the final straw that pushes the economy into a financial stress region, but the build-up starts before that with a low stance. The time gap until too loose a stance translates to systemic financial instability gives policymakers the opportunity to react and release the pressure from financial markets. Consequently, the implementation of well-considered macro-prudential policies, gentler monetary tightening, or the adoption of backstop policies during financial distress (Boissay et al., 2022) could be the key to avert financial crises. However, the efficacy of macro-prudential policies in this context remains an open question and is hard to test with our data as, for most of our sample period, such policies were not part of policymakers' toolkits.

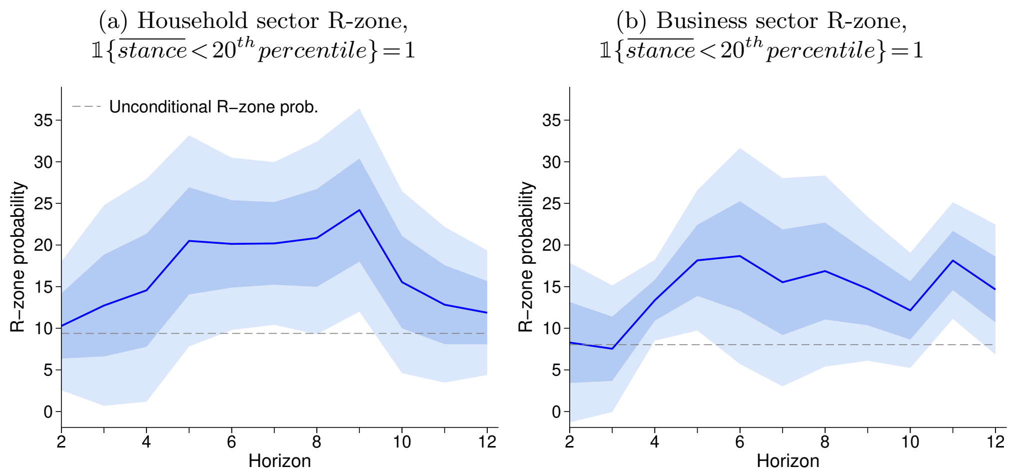

FIGURE 8

The connection between very loose monetary policy and post-WWII R-zones.

Notes: The same notes as in Figure 7 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.

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 often characterized by a reallocation of credit from the tradable to the non-tradable sector, and evidence suggests that only credit expansions to the non-tradable sector are associated with a heightened risk of financial instability (Müller and Verner, 2024). The findings presented in Figure 7 also continue to hold when we enrich the model with decade fixed effects or when we replace global control variables with time fixed effects (Appendix Figure A16). Furthermore, the estimation of a logit model yields similar results (left panels of Appendix Figure A17).

Similar to our earlier exercise with crisis prediction (panel (b) of Figure 4), we can also establish a relation between a simple indicator for *very loose* monetary policy and R-zone risk. This is shown in Figure 8. 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 18.7% 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 et al. (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 9 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

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

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

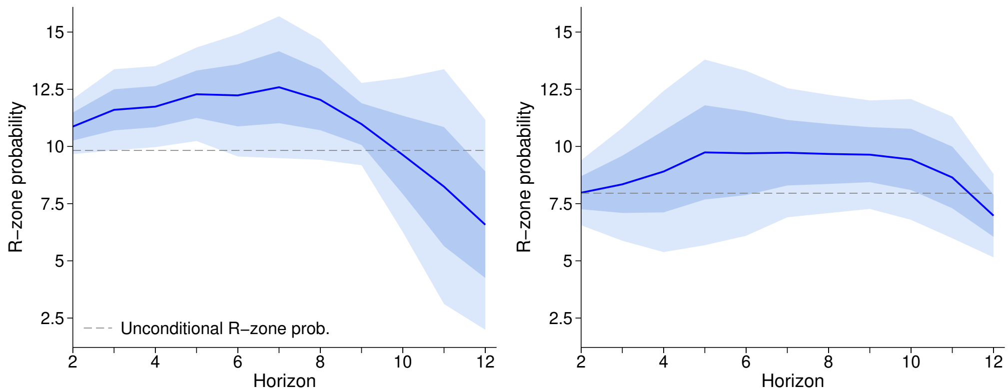

FIGURE 9

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

Notes: The same notes as in Figure 7 apply with one difference; all available data points of the JST Database are now used.

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

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

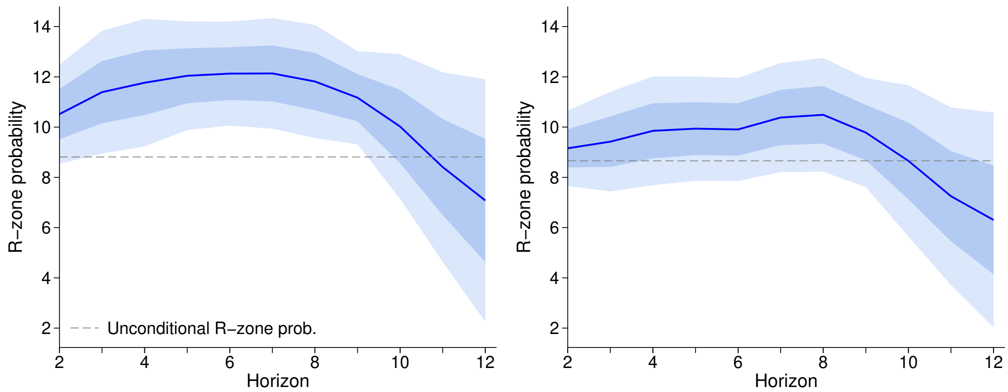

FIGURE 10

The connection between loose monetary policy and "housing finance" R-zones.

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 7 apply.

economic magnitudes of our estimation, a conclusion that also applies to a logistic model (right panels of Appendix Figure A17).

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 et al., 2015a), we refer to this interaction as a Red-zone too. Figure 10 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 7 (a) and 10 (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 9.9% (7.4%) 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—and strengthened—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*

The stance of monetary policy likely responds to economic conditions. Although we try to control for a rich set of observable macrofinancial variables, we may suspect that there remain endogenous components in the fluctuations of stance orthogonal to these controls. In this section, we follow Jordà et al. (2020) who use the trilemma of international finance to construct a historical series of cross-country monetary policy shocks that we can use as an instrument. 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 (Obstfeld and Taylor, 2004). Rather, the forces of international forex arbitrage require the pegging country 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 conditions of the pegging country into account, then the pegging country's monetary policy response to them is exogenous.[13] 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 as

$$z_{i,t} = \begin{cases} k_{i,t}\left(\Delta R_{b(i,t),t} - \Delta\hat{R}_{b(i,t),t}\right) & , \quad \text{if } q_{i,t}=1; \\ 0 & , \quad \text{if } q_{i,t}=0. \end{cases} \tag{4.10}$$

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 et al. (2011) capital mobility indicator (1 if open).[14] $\Delta R_{b(i,t),t}$ is the nominal interest rate change

13. See Jordà et al. (2015a) for a motivation of this identification assumption.

14. The Quinn et al. (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.

in i's base country b in year t and $\Delta\hat{R}_{b(i,t),t}$ are corresponding predicted changes in $\Delta R_{b(i,t),t}$ using lagged base-country predictors.[15]

Summing up, 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. See Jordà et al. (2020) for more details on this instrument, its construction, and a complete overview of the assignment of base countries to pegging countries.

We are interested in the variation in the stance of monetary policy induced by a series of accommodative monetary policy shocks, such as those that hit the European periphery in the 2000s. In the model of Boissay et al. (2022), it is precisely such a prolonged period of expansionary monetary policy shocks that precedes the average financial crisis. To this end, we instrument $\overline{stance}$ with 10 lags of $z_{i,t}$. Recall that we define the stance of monetary policy as the moving-average difference between the real rate and natural rate from year $t-4$ to t (see equation (2.4)). 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.

Appendix Table A1, which presents the first stage regression, confirms this intuition behind our IV choice. A history of medium-term contractionary (expansionary) monetary policy shocks identified by the IV is linked to a significantly tighter (looser) stance of monetary policy in the home economy. Together, the instrumental variables present a strong first stage as indicated by the Kleibergen-Paap (2006) test for weak instruments.[16]

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 not just robust but is actually reinforced. Figure 11 shows that the impulse responses here exhibit a very similar pattern. However, the point estimates are larger, likely because IV mitigates attenuation bias under OLS due to measurement error coming from r^*, as noted earlier.

Our baseline specification indicates that the likelihood of entering a financial crisis 5 to 7 years ahead is higher by 5.8 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 16.4% when $\overline{stance}=-1\%$. The LP-IV estimates suggest that at higher horizons, the medium-term crisis risk more than doubles from 10.5% to over 20%. The inclusion of decade fixed effects does not alter these conclusions, as shown in panel (b).

As in the previous LP-OLS setup (Figure 4), bootstrap confidence bands confirm the significance of the estimates, as shown by the red dashed lines. Furthermore, the results are robust to all the modifications explored earlier. In particular, Appendix Figure

15. 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.

16. 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, Appendix 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.

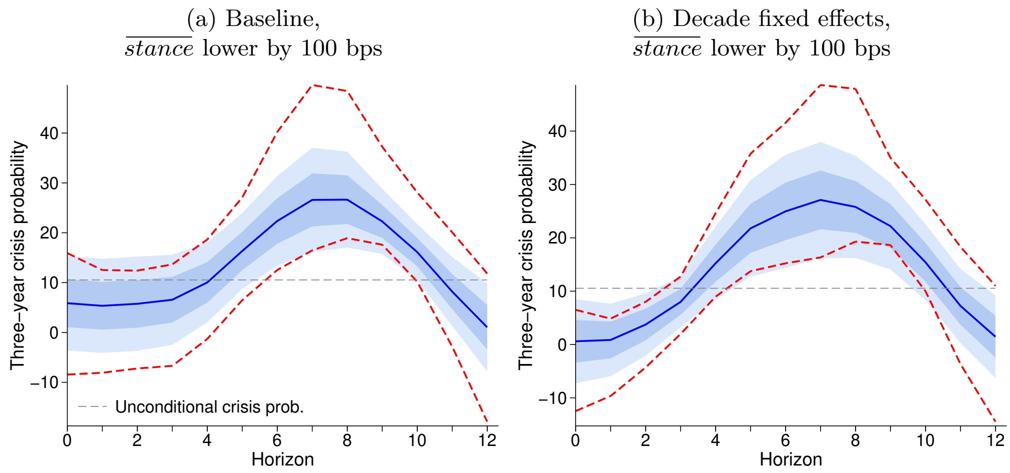

FIGURE 11

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

Notes: In this figure we re-estimate model (2.5) 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 (4.10). The corresponding first stages are presented in columns (1) and (2) of Appendix Table A1. The points show IV estimates of $\{-100\beta^h\}_{h=0}^{12}$. Apart from these modifications, the same notes as in Figure 4 apply.

A18 shows that similar results are obtained when focusing on the post-WWII period. 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 as seen in Appendix Figure A19. Allowing for two lags instead of one in the VAR process in (SE) to estimate r^* and $\overline{stance}$, the ensuing 2SLS estimates are comparable, as confirmed in Appendix Figure A9 (b). Our IV estimates are not significantly affected when we construct our stance variable based on ex-ante real rates, employing the two measures of inflation expectations introduced and explained in Section 3.1. This is illustrated in Appendix Figure A15 (b).

Appendix C provides interpretable bounds on potential spillover effects using the control function approach of Jordà et al. (2020). Appendix Figure C1 shows that these bounds are narrow and closely aligned with our baseline point estimates. Nevertheless, the estimates should be interpreted in light of the untestable nature of the exclusion restriction.

4.3. *Credit and asset market overheating*

In Figure 2 (a) of the Introduction, we illustrated the divergence in the eurozone's monetary policy stance between the core and periphery during the 2000s. This occurred because the central bank focused on the moderately growing core while overlooking the booming periphery. A series of interest rate cuts in the base country can push the interest rates of the pegging country below neutral. We next explore whether such variations in the stance of monetary policy increase the likelihood of entering R-zones, periods of credit and asset market overheating, using our instrumental variable for identification. Figure 12 first shows that the divergence in the monetary policy stance among eurozone

FIGURE 12

Credit and asset prices in the eurozone before the Global Financial Crisis.

Notes: Motivated by the definition of R-zones (equation (3.9)), 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.

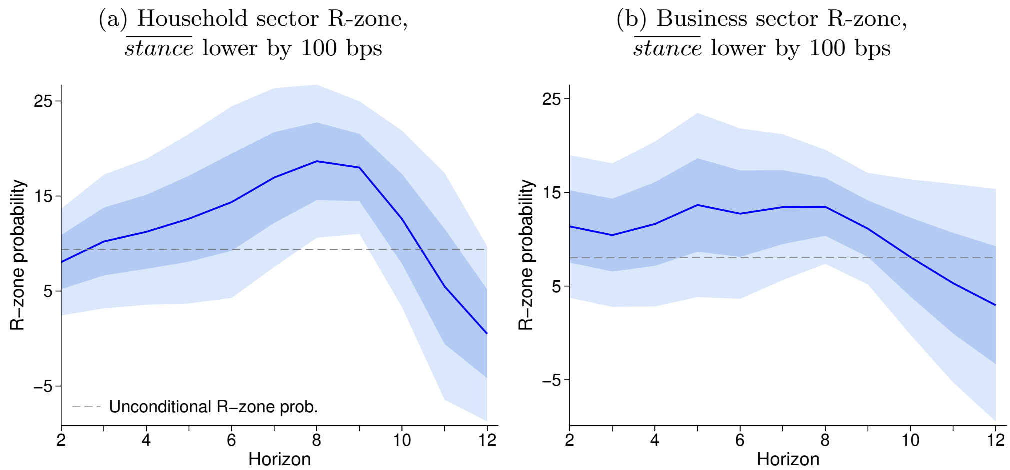

FIGURE 13

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

Notes: The same notes as in Figure 7 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 (4.10).

countries is reflected not only in annual total credit growth, as illustrated in Figure 2 (b), but also in those financial variables used to capture household sector, business sector, and housing finance R-zones. While stock price dynamics were largely synchronized across core and periphery countries before 2008, the credit and house price booms were far more pronounced in the periphery than in the core, i.e., precisely in those countries with "too loose" a stance.

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 reinforce the LP-OLS estimates reported in Figure 7. Figure 13 suggests that a loose stance of monetary policy raises the likelihood of credit and asset market overheating as defined by Greenwood et al. (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 (Appendix Figure A20) and in the full sample (Appendix Figure A21). 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, 2024).

5. A GROWTH-RISK TRADE-OFF

Finally, we note that a loose stance of monetary policy also has potential benefits, as well as costs. Running a "high-pressure economy" may not be a bad thing *per se*.[17] It might, for example, raise the productive capacity of the economy through more efficient utilization of capital and labor and offer greater incentives for spending on research and

17. We would like to note in this context that we verified that as expected, prices rise after a period of loose monetary policy. This is shown in Appendix Figures A22 and A23.

development as well as for launching new, innovative businesses (Okun, 1973; Yellen, 2016).

However, our historical evidence suggests that running such a high-pressure economy may not be sustainable in general. We now argue that potential short-term gains come at considerable medium-term cost in the form of heightened risk of disasters in real economic activity. As noted in the Introduction, this closes the circle between the conduct of monetary policy, financial fragility, and real activity, while connecting to the Growth-at-Risk literature (Adrian et al., 2019, 2022).

We have shown that when monetary policy is too accommodative it 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 (2.5). 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 of left-tail growth risk. Consistent with the non-existent association between a loose stance and elevated short-term crisis risk (Figure 4), the likelihood of entering the left tail of the cross-country real growth distribution stays put in the immediate aftermath of a spell 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 $\overline{stance}$ 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. However, 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).

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)

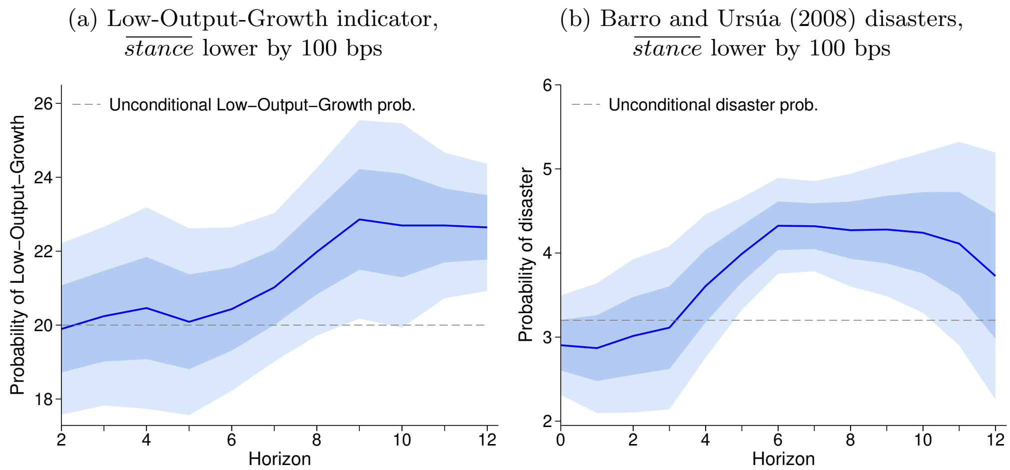

FIGURE 14

The connection between loose monetary policy and the left tail of real growth.

Notes: The figure shows estimates of $-100\beta^h$ of model (2.5) for different horizons h relative to the unconditional full-sample probability that $B_{i,t} = 1$. In panel (a), $B_{i,t} = \textit{Low-Output-Growth}$ 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 $\mathit{ceiling}(1.5 \times h)$ lags.

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. A loose stance over an extended period leads to increased financial fragility several years down the line. Early signs after a loose stance episode are the associated swings in those financial variables that have been identified by the literature as harbingers of financial turmoil, asset prices and credit creation.

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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Supplementary Data

Supplementary data are available at *Review of Economic Studies* online.

Data Availability

The data and code underlying this research is available on Zenodo at `https://doi.org/10.5281/zenodo.19036251`.

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