Essays on Inflation Expectations, Leaning Against the Wind Policy, and Consumer Bankruptcy

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Introduction

Times of severe economic crisis are, fortunately, rare events. Large recessions, such as the Great Depression of the 1930s or the Great Recession during the late 2000s, have detrimental effects on all economic sectors. During such times, the public as well as the academic debate centers around the causes of the crisis and which policies are best suited to alleviate the economic fallout. In its three chapters, this thesis contributes to our understanding of the answers to these questions. It draws on empirical as well as structural macroeconomic results to foster our knowledge of potential causes of, and historical experiences with, deep economic crises. In particular, this thesis presents insights into the German experience during the Great Depression, the link between monetary policy and the likelihood of financial crises, and how competitive pricing of income risk in unsecured credit markets exacerbates business cycle fluctuations.

Chapter 1, titled "Inflation expectations and the recovery from the Great Depression in Germany", is joint work with Volker Daniel and published in *Explorations in Economic History*. It investigates whether changes in inflation expectations can explain the remarkable economic recovery from the Great Depression in Germany. The Great Depression stands out as the most severe economic decline in modern history. For the United States, the implementation of restrictive monetary policy at the time has been identified by Friedman and Schwartz (2008) to have contributed substantially to the depth and duration of the economic decline. At the same time, Jalil and Rua (2016) show that President Roosevelt's credible commitment to inflate the economy in Spring of 1933 lead to changing inflation expectations, which provided the impetus that jump-started the economic recovery. Given that Germany experienced a similarly successful, albeit more steady recovery from the economic mayhem in August of 1932, this paper analyzes changes in expected inflation as a potential explanation for the economic upturn.

Our analysis provides a detailed narrative account of newspaper coverage of inflation over the course of the Great Depression in Germany. In particular, we identify four episodes of increased inflation coverage in the news which we analyze in detail. Our finding is that, while fears of inflation were a prevalent issue at the time in Germany, there was no sustained shift in expected inflation rates in the narrative account. The government and the Reichsbank were always eager to forcefully rule out any inflationary measures whenever the potential of increased inflation loomed ahead. We further corroborate these findings with time series estimates of expected inflation from a FAVAR model, real interest rate regressions, inflation forecasts based on inflation news counts, and historical decompositions. None of the estimates show any change in expected inflation when the recovery phase started. Rather, the initial sparks for the recovery are more likely to have emanated from a boost in confidence as a result of the end of reparations as well as a change in fiscal policy.

Chapter 2, titled "Leaning against the wind and crisis risk", shifts the focus towards the influence of monetary policy on the risk of financial crises. In joint work with Moritz Schularick and Felix Ward we address the question: Should monetary policy lean against credit and asset price booms to safeguard financial stability? In the wake of the Great Recession this question reemerged as booming asset prices and credit were argued to have posed a significant stability risk. This build up in financial markets could potentially have been mitigated by monetary policy raising interest rates above and beyond what was warranted by inflation and unemployment, i.e. by *leaning against the wind* (LAW). Proponents of LAW policies argue that tight monetary policy can lower the risk and severity of financial crises (Borio and Lowe, 2004; Cecchetti, Genberg, Lipsky, and Wadhwani, 2000; Roubini, 2006). On the other side of the argument, critics of LAW policies argue that monetary policy is ineffective in lowering crisis risk (Bernanke and Gertler, 2001; Gilchrist and Leahy, 2002; Svensson, 2017), or outright harmful in the sense that they provoke, rather than prevent, financial crises (Bernanke and Gertler, 2000; Bernanke, 2002). Given this long-standing controversy, what is needed at this point are credible, causal estimates of the effects of LAW policies.

In this chapter, we provide such estimates using a rich dataset on advanced economies since the 19th century. The empirical results are derived from an instrumental variable approach that exploits exogenous variation in country specific short term interest rates as in Jordà et al. (2019). Specifically, under high capital mobility, countries with fixed exchange rates are forced to adjust short term interest rates in accordance with the respective base country. Changes in the base country interest rates are arguably independent of local economic conditions, providing us with exogenous policy rate variation. We find that a 1 percentage point (ppt) policy rate hike during a financial boom increases crisis risk by about 10 ppts within the year. Crisis risk remains elevated for one year thereafter, before subsiding to its long-run average level. At no point in the 5 years following the policy rate increase do we find evidence for a systematic reduction in crisis risk. The empirical evidence thus lends support to some of the worst fears about LAW policy—that it is more likely to trigger crises than prevent them. Furthermore, we show that LAW policies prior to financial crises do not significantly reduce the severity of such crises. Thus, while the crisis trigger effect is a robust feature of the data, there is no evidence for crisis severity reduction effects of LAW policy.

Chapter 3, titled “Destabilizing Effects of Consumer Bankruptcy”, is joint work with Lisa Dähne and studies the link between household income risk, consumer bankruptcy, and business cycle fluctuations. During recessions, charge-off rates on credit card loans rise substantially, a phenomenon which was especially visible during the financial crisis of 2008. At the same time, income risk faced by households also increases in the early stages of a recession. We can now, more than a decade later, recognize very similar patterns as the ongoing Corona epidemic unfolds. With the threat of large scale defaults on loans looming, governments are taking action by granting massive financial aid packages to firms and households. Since unsecured credit is priced based on the likelihood of default, it is important to understand how the link between income risk and consumer bankruptcy shapes individual consumption and savings decisions and aggregate outcomes. Whether the possibility to declare bankruptcy amplifies or attenuates the effects of aggregate shocks is a priori not clear. On one hand, credit becomes ex-ante more expensive as banks charge a premium on debt over the return on safe assets. This limits the access to credit and hence the ability to self-insure. On the other hand, the possibility to declare bankruptcy may act as another way households can smooth out adverse shocks, thus improving the ability to self-insure. This chapter presents a quantitative assessments of these channels.

We first document empirically that increases in household income risk are followed by drops in aggregate output and consumption, as well as tightened lending standards and increasing default rates on unsecured credit. Building on our empirical results, we then embed the theory of consumer bankruptcy in a Heterogeneous Agent New Keynesian (HANK) model with sticky prices and aggregate risk. In the model, households borrow from banks that set the cost of borrowing according to the likelihood of default. An increase in household income risk leads to an increase in the default probability, resulting in higher costs of borrowing and hence tighter borrowing constraints. This subsequently leads to aggravated business cycle fluctuations as default rates increase and households are forced to delever. Our results show that banks adjusting the cost of borrowing in response to income risk shocks accounts for about 30% of the aggregate economic decline. As such, the risk-based pricing by banks makes the economy less resilient overall in response to increases in income risk. This pricing mechanism, however, also amplifies the transmission of monetary policy shocks. A surprise reduction of the interest rate leads to larger increases in consumption and output as future incomes are expected to be higher, making exercising the option to default less attractive. This feeds back into current debt prices, thus loosening borrowing constraints to households. Taken together, our results show that the possibility of strategic defaults introduces an anticipation channel that amplifies the effects of aggregate shocks.

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Chapter 1

Inflation expectations and the recovery from the Great Depression in Germany*

Joint with Volker Daniel

1.1 Introduction

Inflation expectations play a central role in explanations for the recovery from the Great Depression in the United States. Central bankers frequently refer to the historical precedent for this policy prescription: President Roosevelt made a credible commitment to inflate the economy in the Spring of 1933, which is regarded as a regime change that marks the beginning of the successful recovery of the U.S. economy (e.g. Temin and Wigmore (1990), also Eggertsson (2008)).[1]

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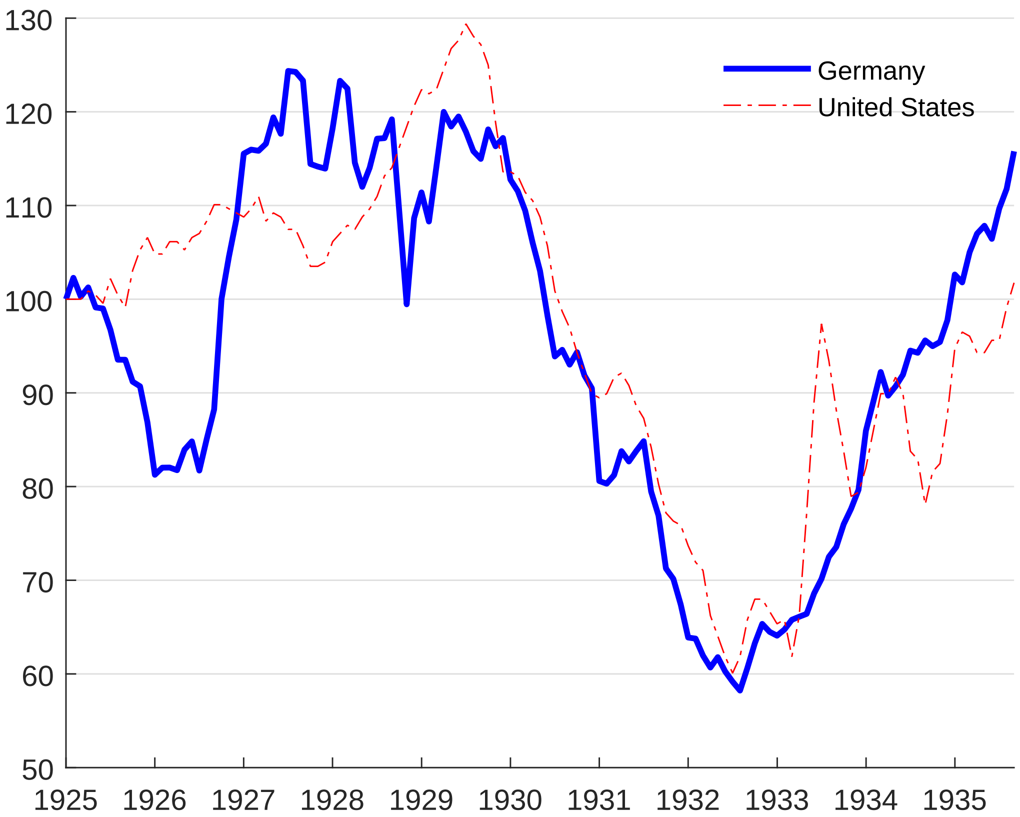
[1] As asserted by Jalil and Rua (2016), the regime change in the United States fueled by inflation expectations initiated the strongest increase in industrial output in a single quarter in U.S. history. Chouliarakis and Gwiazdowski (2016) and Shibamoto and Shizume (2014) showed that inflation expectations also played a pivotal role in Great Britain and Japan escaping the Great Depression,

Figure 1.1: Industrial production in Germany and the United States 1925 to 1935



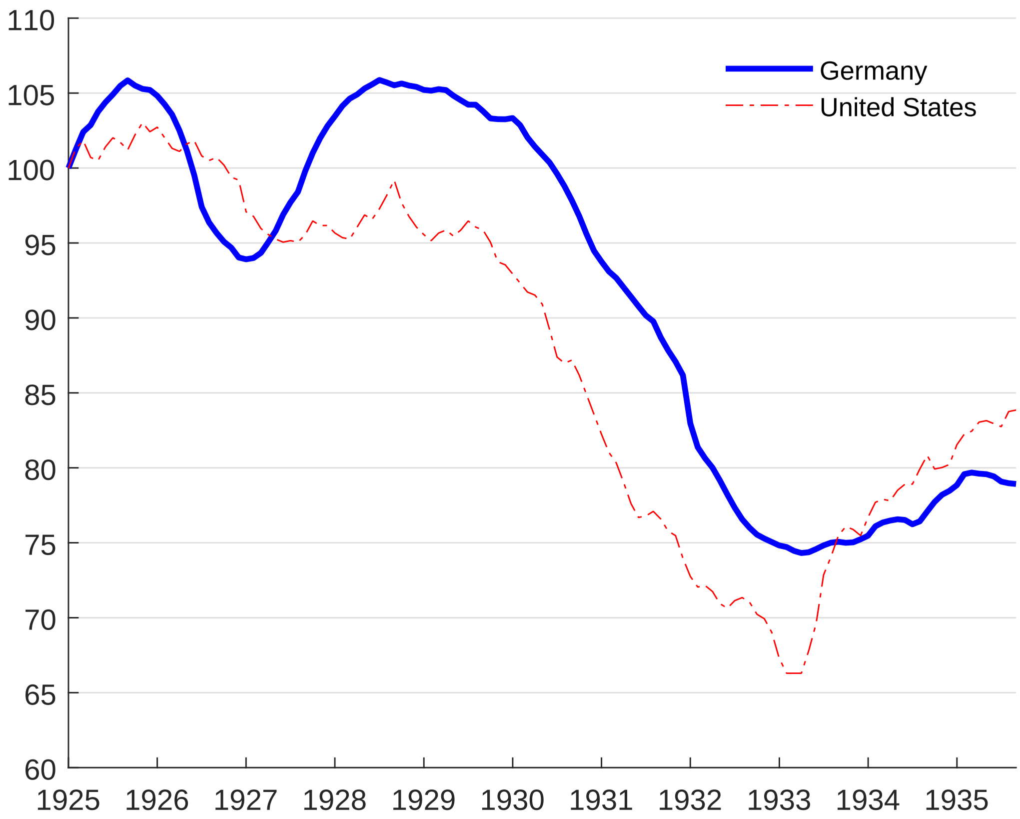
Notes: Industrial production in Germany and the United States, monthly, Jan 1925 = 100. Both series are seasonally adjusted. *Sources:* Wagemann (1935), series III.B.11; Governors of the Federal Reserve Board, Statistical Release G. 17 (INDPRO), (2013).

Figure 1.1 shows the recovery paths for the United States and Germany. U.S. industrial production increased dramatically following the Roosevelt regime shift in the second quarter of 1933. The recovery in Germany that started in 1932 proceeded at roughly the same rate, starting off slowly in the beginning, but at a higher and steady pace thereafter. We observe similar patterns for prices (Figure 1.2). Although the U.S. price index of industrial finished goods jumped upward along with industrial production, the German price index increased at a much slower pace.

In this paper, we investigate whether a shift in inflation expectations initiated the German recovery as occurred in the United States. The debate about the causes for this remarkable economic upturn remains unresolved. Inflation expectations could be a key factor in the recovery of Germany in the 1930s that has not yet been considered. From a theoretical perspective, changing expectations regarding future inflation rates is a crucial driver of production, consumption, and prices in many

respectively.

Figure 1.2: Prices in Germany and the United States 1925 to 1935



Notes: Monthly prices of industrial finished goods in Germany and U.S. index of wholesale prices of finished products, Jan 1925 = 100. Both series are seasonally adjusted. *Sources:* Wagemann (1935), series IX.B.23; NBER Macrohistory database (1997), series 04169.

macro models. In the New Keynesian framework, an increase in expected inflation leads to increased production as consumers substitute consumption over time, and an increased demand for goods leads to an increase in inflation (Galí, 2015). To test whether this channel was operative during the initial phases of the German recovery, we construct inflation expectations using various approaches. First, we conduct a narrative identification of inflation expectations from media articles, following Jalil and Rua (2016). The narrative account is then supplemented with a factor-augmented vector autoregression (FAVAR) model, real interest rate forecasts, and forecasts from inflation news series.

Our central finding is that there is no sustained shift in inflation expectations in Germany at the start of the recovery from the Great Depression in 1932. Although the narrative study identifies occasional fears of inflation on a number of specific dates and events, newspaper articles reveal no sustained regime shift of inflation expectations. Starting with the British exit from the gold standard in September 1931, German newspapers regularly mentioned currency devaluation and inflationary policies as viable policy options. Fears of inflation appeared during discussions about an extensive expansionary policy in January 1932, after the formation of the Papen government in June 1932 and as a response to Adolf Hitler's seizure of power in 1933. However, according to the news account, none of the events caused a sustained change in inflation expectations. Each time the general public thought about inflation, the government and Reichsbank were eager to rule out any price-increasing policies. This finding is in line with what Straumann (2009) argues was the case for several European countries at that time. One explanation for the emergence of fears of inflation among Germans during the Great Depression is their experience with hyperinflation in 1923, a point noted by Borchardt (1985) and Eichengreen (1992), which likely prevented politicians from undertaking inflationary programs similar to those in the U.S.

Our time series estimates also show no indication of a sudden regime shift as an explanation for the recovery. Uniformly across approaches, we observe that during 1932 expected inflation rates remained negative and largely unchanged relative to 1931. Although we do observe occasional evidence that inflation expectations changed from deflation to inflation during the summer of 1932, consistent with the narrative evidence, these expectations did not last. A historical decomposition of industrial production further shows that although changes in inflation expectations were important for explaining the economic decline, they did not provide the initial impetus for the following upswing.[2]

1.2 Literature

The debate about the causes of the remarkable German recovery remains unresolved. Although some scholars have argued for a unique economic upturn under the Nazis initiated by vast investment programs in highways and rearmament (Abelshauser (1999), Overy (1975)), others have emphasized that the starting point occurred before the Hitler dictatorship in 1932 (Buchheim (2003), also Ritschl (2003)). According to this narrative, the Nazi expansionary policies were deemed unnecessary because previous government interventions and the recovery had already begun be-

[2] Many macro models imply that changes in expected inflation rather than the level are relevant for growth. It is, however, empirically plausible that the level of inflation is also important insofar, for instance, as a shift from mild deflation to mild inflation might have outsized effects on the public's economic outlook. Thus, we proceed by examining both changes in inflation and the level of inflation while remaining agnostic regarding which one is more relevant in the interwar German context. Because we find that inflation expectations were little changed, our argument that inflation expectations did not contribute to the recovery holds regardless of whether changes or levels of inflation were relevant. Moreover, changes in expected inflation should have triggered changes in the real rate of interest as a crucial driver of saving and investment decisions. The real rate of interest in Germany remained at very high levels during this period and did not decline at the start of the recovery as it was the case, for instance, in the U.S. in the spring of 1933. We thank the referees for emphasizing this point.

fore the Nazis could specify their economic stance in 1933. Tooze (2006) highlighted the important agreement at the conference of Lausanne in July 1932, which substantially reduced the German debt burden and provided the necessary room to combat the economic slump (also James (1986)). Temin (1990) noted the freezing of wages after 1931 as a monetary explanation for the German recovery. Both arguments resonate the dilemma emphasized by Borchardt (1982): high wages and debt burdens prevented German governments before 1932 from implementing alternatives to the devastating deflation. This brings us to our question: during this heavily deflationary episode, could inflationary expectations have served as a kick starter for the depressed economy? The experience with hyperinflation in the 1920s could have made Germans overly sensitive to news and shocks regarding expected inflation. Inflation expectations may therefore have been a crucial but not yet investigated part of Germany's recovery.

Our paper is closely related to the literature describing narrative evidence for the Great Depression in the U.S., for example, by Jalil and Rua (2016), Nelson (1991) and Romer and Romer (2013). Quantitative forecasting methods usually have a backward-looking perspective and depend on the choice of variables, whereas narrative evidence incorporates ideas and considerations at any time point that may be independent of reflections on the past and an available set of data. For this reason, they detect a regime shift that time series approaches fail to identify (see Romer (2013), Sargent (1982), Temin and Wigmore (1990)). The narrative approach can identify the current knowledge of the general public and which information could be considered relevant to changes in inflation expectations. Identifying the actual sources is crucial, because a shift toward inflation expectations should not be a purely statistical outcome; it should be experienced by contemporaries due to an abrupt policy change, an event, or a shock to be identified as a regime change (Jalil and Rua, 2016).

The use of direct time series forecasts using factor models has gained considerable popularity for forecasting economic time series. Boivin and Ng (2005) and Eickmeier and Ziegler (2008) evaluated forecasting methods and found that factor-based forecasts perform well in practice and tended to perform better than simpler small-scale models. Therefore, a factor model is a natural candidate for our research question. A key advantage of factor models is that they allow us to incorporate much information into the analysis while keeping the estimation procedure tractable. Furthermore, because factor models are estimated on the basis of the comovement between the time series in the dataset, a significant expansionary shift visible in many time series would be picked up by the factors, and this should provide information regarding future inflation rates if the underlying cause of such an expansionary shift was because of changes in expected inflation rates. One limitation with time series approaches similar to the one used in this study is that they do not identify the correct expectations in the event of a sudden regime shift. Therefore, we carefully discuss possible shifts in expectations in the narrative account.

One prominent method to estimate inflation expectations, which uses the Fisher equation that relates expected real interest rates to nominal interest rates and expected inflation, was first proposed by Mishkin (1981). He showed that rational expectations imply that inflation expectations can be inferred from realized real interest rates. For the Great Depression in the U.S., Cecchetti (1992) and Romer (1992) applied this approach and found a shift in inflation expectations in 1933 at the beginning of the recovery, at the same time that Temin and Wigmore (1990) detected a regime change toward a more expansionary macroeconomic policy. For Germany, Voth (1999) applied the Mishkin method during the interwar period and detected inflation uncertainty and fears of inflation in 1931 and 1932, but his results did not indicate a shift in inflation expectations. One potential problem with the German interwar financial market data is that the nominal rate or bond yield series reflect liquidity and default risk. This makes disentangling expected inflation from risk or liquidity premia difficult, a point that Voth (1999) similarly recognized in his work.[3] We, therefore, also adopt the approach by Binder (2016), who proposed to regress future inflation rates on a narrative news measure. This measure included scaled counts of articles containing inflationary and deflationary terms. The idea is that the degree of news coverage about inflation should be positively correlated to inflation expectations, which allows us to directly map our narrative evidence into inflation expectations.

1.3 Narrative account

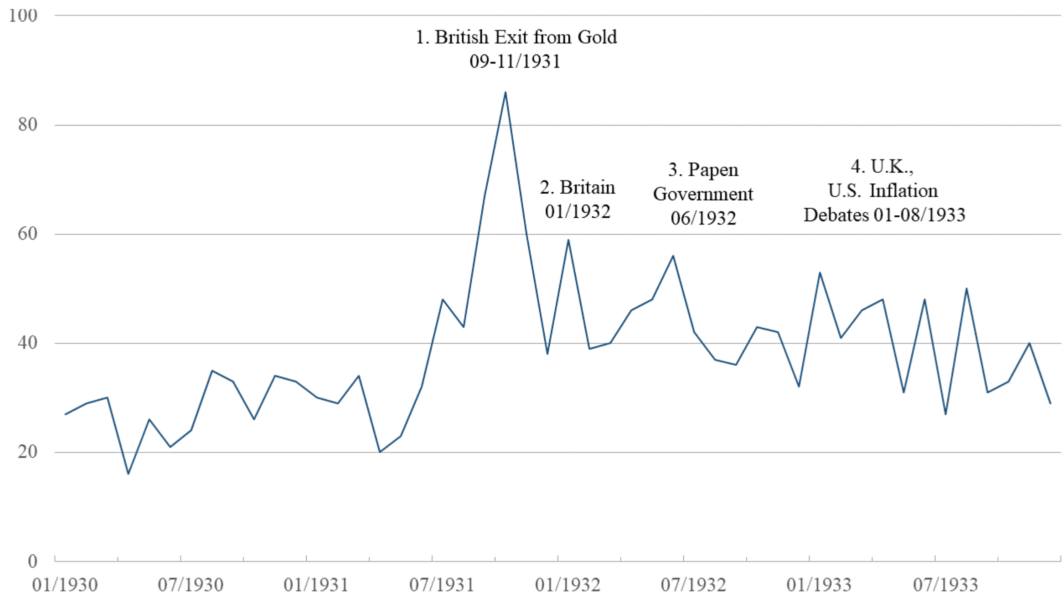
To discover a possible shift in inflation expectations in Germany at the start of the recovery from the Great Depression, we follow Jalil and Rua (2016) and first provide a general overview of newspaper coverage regarding inflation through a word search of newspaper articles. Next, we conduct a detailed narrative study of media sources over a two-year period.

1.3.1 General overview of inflationary news coverage 1930 to 1933

As our main source of media coverage, we study *Vossische Zeitung*, one of Germany's national newspapers of record during the Weimar Republic. *Vossische Zeitung* had a daily circulation of approximately 68,000 in 1931 and covered the main events and

[3] Another reason for a likely distortion was that interest rates and many prices at the time were controlled by effective cartels and government regulations. We thank Carsten Burhop, Werner Plumpe, Mark Spoerer and Jochen Streb for this very important indication.

Figure 1.3: Inflation mentions in *Vossische Zeitung* 1930-1934



Notes: The monthly frequency of articles that contain terms related to inflation in *Vossische Zeitung*. We also indicate prevalent topics in the news in four periods of increased news coverage.

debates of that time (Binkowski and Schottenloher (1985), Deutsches Institut für Zeitungskunde (1932)).

In the online database provided by the De Gruyter publishing house (De Gruyter, 2010), we searched all issues of the newspaper from 1930 to 1933 and counted the number of articles that contained the German terms for "inflation". Specifically, we used a combination of the terms, word stems, and abbreviations of the German words "Inflation" (inflation), "Teuerung" (price increase), and "Reflation" (reflation).[4] If there had been a sudden shift in inflation expectations, there should have been an increased public interest likely reflected by an increased number of articles on inflation in one of the most relevant newspapers at that time.

We depict the resulting series in Figure 1.3. Before the summer of 1931, fewer than 30 articles per month on average mentioned inflation; in September and October, the mentioning of inflationary terms increased threefold to 86 articles. This result signifies that more than two articles per day on average mentioned inflation. After the fall of 1931, inflation coverage occurred to a lesser extent in *Vossische Zeitung*, although appearances remained at a higher average than before, with some months clearly above the number of articles per month before fall 1931. This was the case for November 1931 and for January and June 1932. In 1933, inflation cov-

[4] We also considered terms like "Geldentwertung" (debasement), "Preissteigerung" (price increase), "Preisanhebung" (price lift), "Preiserhöhung" (price increase), and "Preisheraufsetzung" (price increase, markup).

Table 1.1: Months and events with increased inflation mentions in *Vossische Zeitung*

Spike	Month	# Articles	Event/Topic 1	Event/Topic 2
1	September 1931	68	Hyperinflation in retrospect (24)	British exit from gold (20)
	October 1931	86	Germany's economic policy (37)	British exit from gold debate (20)
	November 1931	63	Exit from gold policy debate (24)	Hyperinflation in retrospect (20)
2	January 1932	59	Hyperinflation in retrospect (30)	Exit from gold policy debate (15, in part Wagemann's Plan)
3	June 1932	56	Papen government (21)	Hyperinflation in retrospect (17)
4	January 1933	53	Inflation debate globally (24, in part Hitler)	Hyperinflation in retrospect (19)
	April 1933	49	Inflation debate U.S (37)	Hyperinflation in retrospect (9)
	June 1933	50	Inflation debate U.S (20)	Hyperinflation in retrospect (18)
	August 1933	50	Inflation debate globally (20)	Hyperinflation in retrospect (18)

Notes: The first column indicates the assignment of each month to one of the four spikes indicated in Figure 1.3. Columns 2 and 3 indicate the selected months and the number of articles per month that mention inflation. The last two columns show the two main events / topics involving inflation. Numbers in brackets indicate the number of articles per topic, and in some cases a certain aspect of the event or debate.

erage was lower but still had spikes above the average in January, April, June, and August.

We compared our article counts for Germany to the United States by using articles from *The New York Times* (ProQuest, 2004). In Figure 1.C.1 of Appendix 1.C.1, the word counts from *Vossische Zeitung* in each month on a daily basis reveal that no spike was comparable to the case of the U.S. in April 1933, where article counts regarding the term inflation in U.S. newspapers increased approximately tenfold. Although this indication suggests that news coverage in Germany did not address the topic of inflation as much as the regime shift occurring under the new Roosevelt administration, it is possible that Figure 1.C.1 in the Appendix conceals sudden changes in the German news account. Accordingly, in Figure 1.3, we identify four spikes (or periods of spikes) of increased inflation mention: one outstanding increase in fall 1931 and further spikes in January 1932, June 1932, and January to August 1933. Each spike could indicate a shift in inflation expectations in the news and we therefore discuss them in greater detail in Table 1.1.[5]

To investigate whether any of these spikes corresponded to an inflationary shock, we checked all articles containing inflation during the four mentioned time periods.

[5] As a robustness check, we collected other available newspaper data from searchable databases of newspapers from the 1930s. The remarkable spike in October 1931 is persistent across newspapers from different political backgrounds and locations, and we illustrate this phenomenon in an additional plot of five newspaper series in Appendix 1.C.2.

Table 1.1 indicates the events or topics that corresponded to inflationary words ordered chronologically by the four spikes (column 1) and months of occurrence (column 2). In each month, we examined the context in which inflationary terms were mentioned by reading the relevant articles. We classified articles with reference to a specific current event as dealing with such event (columns 3 and 4). In many cases, inflation was exclusively mentioned in reference to the hyperinflation of 1923 with no relation to the present. In that case, we classified it as "Hyperinflation in retrospect". If an article compared actual events with past periods, we classified the article as referring to the actual event. The few articles not counted as part of either an actual event or the hyperinflation of 1923 referred to a wide range of topics from literature, culture, and sports and were irrelevant to our research question. In columns 4 and 5, we further indicate in brackets the number of articles that we related to the two relevant topics. These two inflationary topics or events were prevalent in many articles that mentioned inflation in the designated months. Notably, more than two-thirds of the articles in our sample each month (comparing the numbers in brackets with column 3) referred to these two subjects. In addition to mentions of the hyperinflation of 1923, we identified references to at least one actual political event or debate in each month. We have indicated these events in Figure 1.3.

In September 1931, inflationary news predominantly covered Britain abandoning the gold standard and related this event to expected inflationary tendencies in Britain. In the following months, the discussion about Britain remained prominently in the news at least until January 1932, as indicated by the topics in Table 1.1. As a consequence of Britain abandoning the gold standard, the news account in *Vossische Zeitung* reveals that after the fall of 1931, inflationary policies and a devaluation of the Reichsmark were also discussed in Germany. German economic conditions were regularly compared with those of other countries such as Britain and also to Germany's experience with hyperinflation. Importantly, the news account also shows that such considerations were regularly opposed by the German government and central bank.

In January 1932, during the ongoing inflationary debates in Britain and the U.S., a plan for a more expansionary economic policy proposed by the head of the statistical office, Ernst Wagemann, received some media attention. In June 1932, the new Papen government was allegedly expected to implement inflationary policies. In January 1933, this was the case when Hitler seized power. Inflationary policies in the U.S. were in the focus of the now controlled media for the remainder of 1933. The word counts visualized in Figure 1.3 are therefore remarkably in line with Borchardt's (1985) notion that fears of inflation prevailed after political events associated with currency devaluation, credit expansion, or expected government deficits.[6]

1.3.2 A detailed narrative study of media sources September 1931 to August 1933

Next, we conduct a chronological investigation of the relevant political events and debates regarding a possible shift in inflation expectations through a careful narrative study of media sources. For this study, we read weekly issues of the economic periodical *Der Deutsche Volkswirt (1931/1933)* over a two-year period from September 1931 to August 1933. The weekly frequency permitted a careful reading of articles over a prolonged time period. We divided our analysis according to the spikes in the news coverage discussed in Section 1.3.1 and included potential inflationary events during other periods. As one of the leading and influential periodicals in its field at that time (Röpke, 1933), *Der Deutsche Volkswirt* had up to 6000 subscriptions and provided a summary of political events and the state of the economy (Sattler, 1982).

For relevant dates, we considered two major daily newspapers: *Vossische Zeitung (1931/1933)* and *Berliner Lokal-Anzeiger (1931/1933)*. We read both newspapers one week before and after such dates to observe responses and to identify diverse opinions regarding the events. While *Der Deutsche Volkswirt* and *Vossische Zeitung* were considered economically liberal and politically centrist, *Berliner Lokal-Anzeiger* supported more nationalist, conservative positions. With a daily circulation of approximately 200,000 in 1932, it was an influential news source of the political right. As a robustness check, we considered further news sources that spanned the entire political spectrum and contemporary scientific publications. For instance, we assessed the coverage of relevant events in the social-democratic newspaper *Vorwärts (1930/1933)*, the catholic-centrist *Germania*, and the weekly reports of the Institute for Business Cycle Studies.

[6] As Eichengreen (1992) noted, events that reminded Germans of the hyperinflation in 1923. In Appendix 1.C.3, we plot an additional word count that dates back to 1918. This Figure 1.C.3 indicates that our measure of inflationary terms reveals inflationary events and fears of inflation, as described by Borchardt (1985), before the 1930s. A potential concern is that articles that mention inflation actually discussed its opposite: the ongoing deflation and price declines present during the early 1930s. As a robustness check, we conducted word searches of terms related to deflation. We also assessed if price reductions were discussed in the articles mentioning inflation. We verified that this was not the case and the ongoing deflation was not considered a problem in the news account. We illustrate this finding by contrasting scaled word counts of "inflationary" and "deflationary" terms in Appendix 1.C.4.

1.3.2.1 September to December 1931 - Britain abandons the gold standard and inflation debates in the Reichstag

The first potentially inflationary situation occurred in September 1931. After Britain left the gold standard on September 21, on several occasions, the German government and Reichsbank announced that they were "not considering devaluing the Reichsmark for the possible loss of currency stability and inflationary consequences" (*Vossische Zeitung*, September 25-30, 1931). *Der Deutsche Volkswirt* (September 25, 1931) argued that similar measures in Germany could cause inflation because of its experience with hyperinflation in 1923 and claimed that fears of inflation were already fueled by right-wing calls for more autarky, that is, to abandon the gold standard and establish an unconvertible "interior currency" in Germany. This claim was not far-flung, considering that some industry groups and numerous conservative figures somewhat openly demanded deflation to end (*Berliner Lokal-Anzeiger*, September 20, 1931, *Vorwärts*, September 19, 1931).

Two weeks later, the political right announced a vote of no-confidence against the Brüning government for the following parliamentary sessions (*Vossische Zeitung*, October 12, 1931). For the remainder of the parliamentary week, inflation became a persistent topic during the important speeches. Speakers from the moderate left to the center-right defended the government's policy of economic stability while denouncing the nationalists for allegedly turning Germany into inflationary chaos should their no-confidence vote succeed. On October 16, - the day of the vote - *Der Deutsche Volkswirt* declared that the fall of the government as threatened by the "right-wing-opposition, raised fears of another inflation" that already articulated in "panic buying and a retention of the sellers" (see also *Vossische Zeitung*, October 13-15, 1931).

After Brüning's victory in the confidence vote, *Der Deutsche Volkswirt* (October 23, 1931) summarized the dramatic events of the preceding week as a "commotion of mistrust in the banks and toward the stability of the currency". It recognized an increase in acquirements and an expansion of retail sales due to panic buying in the preceding week (October 30, 1931). In November, news regarding inflation abated, and no "inflationary events" were reported in the media. Possible inflationary tendencies in England and other countries that had abandoned the gold standard were frequently discussed, and potential currency experiments in Germany were regularly opposed by politicians and interest groups (see *Vossische Zeitung*, November 1, 6, 10, 24 and 28, 1931). Reasons for the strong opposition to inflationary tendencies can be found in a statement by Hans Luther, head of the central bank, from November 24, 1931:

"Beyond what is happening already there is nothing that can be done through either currency or credit policy to provide an impulse for economic recovery. [...] The decision by the German government and the Reichsbank to not let the Reichsmark float on September 20 was the necessary conclusion given German indebtedness and inflation experiences by the German people."

The German experience with hyperinflation thus seemed to be a strong factor that determined policymaking. This topic returned two months later when Ernst Wagemann, head of the German statistical office, asserted his proposal to end the Depression. One consequence of the currency devaluations in other countries was that economic discussions in Germany shifted to the decrease in prices and wages to maintain Germany's competitiveness in world markets, which basically constituted the opposite of inflation expectations. The result was the emergency decree of December 8, which forcibly cut prices, rents, wages, and interest rates as of January 1, 1932.[7] First price reductions as a reaction to the decree came within a few days (*Vossische Zeitung* December 18, 1931). During the debate about the decree and in the following weeks, news coverage revealed no indication of a potential increase in prices in the future due to the drastic cut in the present.

1.3.2.2 January to May 1932 - Expansionary policies are ruled out

In January 1932, Ernst Wagemann, head of the Institute for Business Cycle Research and president of the statistical office of the Reich, published a plan to counter the Depression in Germany. The proposal comprised work programs financed by a moderate credit expansion that considered mild inflation (*Der Deutsche Volkswirt*, January 22, 1932, *Berliner Lokal-Anzeiger*, January 20, 1932, also Wagemann (1932)). Some observers indicated positive examples from other countries, and the industry-leaning press hoped for a reflation of the devalued economy. *Der Deutsche Volkswirt* (January 29, 1932) - and as it argued "all experts" - declined Wagemann's idea as either too small and therefore useless or devastatingly inflationary. The economist Carl Landauer warned in an article published on January 22, 1932, that Wagemann's plan could result in "the same inflationary export premium that we know from the years 1921-1923." The editor of *Der Deutsche Volkswirt*, Gustav Stolper concluded his article from February 12, 1932, with the words: "But Wagemann's reform plan would be the safest option to instantly tear down the ramparts that surround the German banking- and currency system; the ramparts that protect it against the rush of inflationary tendencies, and secure its indispensable remaining trust."

Under the impression of this continuing discussion, the German government announced on January 30, 1932, that it would not consider currency experiments or a change to the Reichsbank law (*Vossische Zeitung*, January 30, 1932). Following this virtual refusal of the Wagemann plan, no news was published thereafter that men-

[7] We describe the emergency decree in detail in Appendix 1.A.

tioned inflationary fears or concrete expansionary policies; therefore, no permanent shift in inflation expectations or a “reflation” of the depressed economy occurred.

In the following months, further public work programs drew support from agricultural and industrial interest groups, banks, and trade unions (e.g., *Der Deutsche Volkswirt*, February 19, 1932, *Vossische Zeitung*, January 30, 1932). Hitler’s national socialists put credit-financed work programs into their 1932 party manifesto, and its electoral success may in part be attributed to this unique feature. A common feature of all the proponents of work programs was emphasizing that their plans were too small to be inflationary (e.g., *Vossische Zeitung*, April 14, 1932, also Borchardt (1985)).

1.3.2.3 June to December 1932 - The Papen government

After Brüning’s centrist minority government ended on May 30, stock markets rose, bond prices fell, and cash withdrawals were reported. *Vossische Zeitung* (June 1-2, 1932) noted that “certain circles advocating currency experiments observed their time coming” while “scared capitalists engaged in stocks as a safeguard”. *Der Deutsche Volkswirt* (June 3, 1932) interpreted the situation that “under the prospect of the next, more right-wing government, the general public takes flight into real assets” and concluded: “The events of the last days are object teaching how credit expansion and currency experiments in Germany would take effect. We only hope, the next government will understand.” By contrast, *Berliner Lokal-Anzeiger* (June 1, 1932) suspected “certain circles” of engaging in the stock market to “stage a flight into real assets”. Additionally, “the left-wing press gave the impression that new inflation was about to come.” Both were intended to “increase nervousness”, “panic” and a “catastrophe mood” in Germany. Regardless of whether this interpretation was correct, the media coverage regarding the danger of inflation certainly induced or reflected fears in parts of the public.

When it was announced that Franz von Papen would form the next government, the “dark nightmares” of inflationary measures were suddenly possible (*Der Deutsche Volkswirt* June 3, 1932). The new chancellor appeared willing to form a majority in the Reichstag with German nationalists and potentially even backed by Hitler’s national socialists, which seemed to favor inflationary measures. The media speculated that Reichsbank president Hans Luther could resign for his inconsistent views on stable economic policies (*Vossische Zeitung* and also *Vorwärts*, both on June 2, 1932).

After Papen’s appointment, however, he immediately met with the central bank president and made clear that Luther would continue, and they both publicly declared that “any currency or credit experiments, that could possibly endanger the value of the currency were out of question” (*Vossische Zeitung* June 3, 1932). This action seems to have been sufficient to stop the flight into real assets. The "severe psychosis faded after the new government left no doubt about currency stability" (*Der Deutsche Volkswirt*, June 10, 1932). In the following weeks, no indications of lasting fears of inflation, an end of deflation, or expected price increases appeared in the news accounts.

In July 1932, the Lausanne Conference reached an agreement, namely, a cut in reparations by 90% and, hence, new financial scope for action to fight the Depression. The quasi-end of reparations as negotiated in Lausanne could have provided the conditions for an economic upturn in Germany, although many observers (especially from the right-wing press) did not praise this remarkable success (*Der Deutsche Volkswirt*, July 15, 1932). The agreement implied a gain in sovereignty: implementing Lausanne would mean the end of international supervision of the Reichsbank, which was the policy under the Young Plan, and room to maneuver in monetary policy, a point that Hans Luther was pessimistic about. The president of the Reichsbank warned that Germany should by no means devalue from gold as a consequence of the newly gained freedom because such measures could lead to high rates of inflation. In the following weeks, bond prices increased, possibly because of an expectation of quick discount rate cuts or a change in the Reichsbank law (*Vossische Zeitung* July 9, 1932); however, no news indicated expected inflation.

In late August 1932, under the impression of growing political violence and election campaigns, chancellor Papen announced an expansionary economic agenda. Although work programs had been discussed in the preceding months, *Der Deutsche Volkswirt* (September 2, 1932) stated that the "timing and announcement of the Papen program was psychologically quite effective" and that market participants were initially surprised by its dimension (also *Vossische Zeitung*, August 29, 1932). The plan comprised moderately sized work programs, tax reductions for hiring, and subsidies for building repairs and was financed by central bank loans and wage cuts.

The implemented measures could have potentially fueled inflationary tendencies or reduced deflation (as mentioned at least once in *Vossische Zeitung*, August 28, 1932). However, the media did not expect inflationary tendencies of the measures. The social-democratic newspaper *Vorwärts* highlighted the stock market surges, but it explicitly did not link them to fears of inflation: "stocks of firms that would possibly benefit most from wage cuts proposed by the program gained most". Hence, if the program had positive effects on the economy, contemporary observers expected them through wage reductions, a small credit expansion, wage subsidies, and incentives (e.g. Carl Landauer in *Der Deutsche Volkswirt* October 26, 1932). None of the measures could be related to price increases or to end deflation. The government underscored that it would refrain from currency experiments (*Der Deutsche Volkswirt*, September 9, 1932), whereas the central bank had declared earlier that it was funding projects only if no inflationary policies were implemented (*Vossische Zeitung*, August 24, 1932).

The narrative evidence, therefore, provides no indications for a lasting shift in inflation expectations as a result of this program. In December 1932, under the short-lived Schleicher government, further work programs showed a comparable reaction without expected price increases mentioned in the media.

1.3.2.4 January to August 1933 - The Hitler administration and U.S. inflation

Hitler's seizure of power on January 30, 1933, shocked most German public - politically but also economically. The Nazis had favored a costly large-scale extension of work programs. Compared to other politicians, to reach their goals, national-socialist speakers appeared sufficiently willing and aggressive to use drastic measures: autarky, currency devaluation, a large deficit, and the violation of international treaties. Gustav Stolper interpreted in *Der Deutsche Volkswirt* (February 2, 1933) the resulting dramatic fall in bond prices and increases in stock prices as motivated by fears of inflation: "The economy is paralyzed again by uncertainty about what will come, despite assertions that economic and currency experiments would be ruled out."

Within days, possible fears of currency experiments were no longer mentioned in the media. On the one hand, the government repeatedly announced that it had no such plans and emphasized the importance of absolute security for the German people and the economy; on the other hand, pressure on the free press intensified and economic opinion articles were published less frequently. In one of those articles, Hans Luther warned, in February, that the fall of the international constraints under the Lausanne treaty could be exploited to put the central bank under political influence with unpredictable consequences for financial stability (*Der Deutsche Volkswirt* February 24, 1933). *Vossische Zeitung* (March 10, 1933) discussed the possibility that as part of the consolidation of powers under the new government (Gleichschaltung), the Reichsbank might be the next institution to become disempowered and highlighted Luther's achievement of currency stability. Nonetheless, Luther's resignation as central bank president one week later was not being debated critically. Several newspapers printed Luther's open farewell letter, in which he mentioned his relentless stance in favor of central bank independence and emphasized that Hitler himself had assured him that no currency experiments were planned. His successor, Hjalmar Schacht, declared currency stability a central objective of the Reichsbank (*Vossische Zeitung* April 7, 1933) and continued to pursue an orthodox policy (see James, 1993).

Despite possible expectations of drastic policy changes under the Hitler government, we find no shift toward inflationary policies or expected inflation in the first months of the new government. The Hitler dictatorship clearly did not conduct an openly inflationary policy. Notably, Hitler himself opposed inflationary policies and regarded the power of the state (in the shape of stormtroopers and concentration camps) as a decisive safeguard against it (James, 1986). In a unique example of this power, on May 16, the Munich police accused 200 small businessmen of raising prices and detained them in the Dachau concentration camp. The incident was made public as a cautionary tale, and the public was requested to report imitators (*Vossische Zeitung* May 21-22, 1933, also *Völkischer Beobachter* May 20,1933; see Domröse (1974)).[8] One factor demonstrating the resoluteness of the regime in terms of price increases may be because of the implementation of a general freeze on pay increases shortly beforehand. The freezing of wages meant that price increases became less likely in the short and medium term and made them even more unpopular among the working population.

Between April and August, the news coverage mentioning inflation focused mainly on the U.S., where the Roosevelt administration had been pursuing its inflationary program since April 1933 (*Vossische Zeitung*, April 19-22, 1933). The measures taken in the U.S., especially the devaluation of the U.S. dollar, were welcomed in the German news because of the high indebtedness of the German economy to the U.S. and implied a real reduction of the German debt burden. A devaluation of the Reichsmark would have countered this effect and was not considered: *Der Deutsche Volkswirt* (for instance April 28, May 5, June 2-30, July 14, August 4, 1933) made the case that Germany and the U.S. were different and an inflationary program similar to the U.S. would either be useless or harmful. Therefore, no policy measures were implemented with the expressed intention of raising price levels, and the government and Reichsbank repeatedly ruled out inflationary policies (*Vossische Zeitung*, April 4, June 2, August 10, August 26, 1933). The few incidents in which government intervention resulted in increased prices of particular goods were carefully discussed and the restricted scope of such increases was emphasized. For example, this was the case for minimum prices on animal fats to support suffering farmers and a revision of the cartel law (*Vossische Zeitung*, May 22-23, July 15, July 18, August 29, 1933).

In summary, we observe no regime shift to inflation expectations under the Nazis before September 1933. We likewise detected no indication for the expressed aim to end deflation. Wage controls certainly reduced the acceptability of price increases among the working population.[9]

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8 We thank Harold James for emphasizing this important incident.

9 Because of the likely bias of the media account due to suppression and government control, we verified that there were no inflation expectations in the first months of the Third Reich by examining further sources: the briefings to the press of the Reich Ministry of Public Enlightenment and Propaganda (Reichsministerium für Volksaufklärung und Propaganda) and the reports of the

1.4 Time series evidence

In this section, we employ a factor model to estimate inflation expectations. Factor models have gained considerable popularity for forecasting economic time series. Boivin and Ng (2005) evaluated forecasting methods and found that for prices and one-month-forecast horizons, factor-based forecasts performed better than simple AR(1) forecasts. Bernanke and Boivin (2003) showed that factor models produce forecasts of similar accuracy as the Federal Reserve Greenbook forecasts. Eickmeier and Ziegler (2008) also showed that factor models perform very well than simpler benchmark models or small-scaled models. Therefore, a factor model is appropriate to answer our research question. In addition, we augment this approach with real interest rate regressions as in Mishkin (1981) and quantitative news estimates proposed by Binder (2016).

1.4.1 The empirical model

The factor model we use relates a large number of time series Y to a small number of common but unobserved factors f. The dynamics of these factors are described by a vector-autoregression (VAR) process. Formally, the model is given by

$$Y_t = \Lambda f_t + e_t \tag{1.1}$$

$$f_t = B_1 f_{t-1} + \cdots + B_L f_{t-L} + v_t \tag{1.2}$$

$$e_t \sim \mathcal{N}(0, \Omega), \quad v_t \sim \mathcal{N}(0, \Sigma) \tag{1.3}$$

In Equation (1.1), Y_t is a 109×1 vector of observed variables, f_t is a 3×1 vector of common latent factors , Λ is the corresponding 109×3 matrix of factor loadings, and e_t is a 109×1 vector of idiosyncratic errors. We assume that e_t and v_t are uncorrelated and that Ω is a diagonal matrix. Equation 2 specifies the dynamics of the factors as a VAR with corresponding 3×3 coefficient matrices. Because we use monthly data, we set the lag length to $L = 12$, which is the most commonly used lag length for monthly VAR models. As in any factor model, we need to address the issue that the common factors and loadings are not separately identified. We resolve this issue by following common practice and restricting the upper 3×3 block of Λ to be the identity matrix. The model is estimated using Bayesian methods. The specification of the prior distributions for the parameters follows Ritschl and Sarferaz (2014). We describe the prior distributions and the estimation procedure in detail in Appendices 1.B.1 and 1.B.2 (see also Kadiyala and Karlsson (1997) and

secret state police (Gestapo), available in the German federal archives (Bundesarchiv in Berlin-Lichterfelde) and the Prussian Privy State Archives of the Prussian Cultural Heritage Foundation (Geheimes Staatsarchiv Preußischer Kulturbesitz, Berlin). We thank Mark Spoerer for emphasizing this point.

Kim and Nelson (1999)). With this model, we aim to produce h-step out-of-sample forecasts of inflation rates, conditional on information available at some point in time T. This is easily done using the state-space form of the model in Equations (1.1) and (1.2), which implies the following expressions for the forecasts

$$F_T = [f_T' \; f_{T-1}' \dots f_{T-L}']' \tag{1.4}$$

$$f_{T+h|T} = JB^h F_T + J \sum_{j=0}^{h-1} B^j \tilde{V}_{T+j} \tag{1.5}$$

$$Y_{T+h|T} = \Lambda f_{T+h|T} + \tilde{e}_{T+h} \tag{1.6}$$

Equation (1.4) combines the relevant factor values into a single column vector, according to the lag length in the factor VAR. Equation (1.5) then uses the VAR system to produce a forecast of the factors. Matrix J refers to the matrix that selects the first three rows of the companion form forecasts of the common factors, and matrix B is the companion form coefficient matrix. Lastly, Equation (1.6) uses the factor forecasts to produce the h-step ahead forecast of the panel. Note that in Equations (1.5) and (1.6), we add random errors drawn from their respective posterior distributions to accurately reflect the uncertainty associated with the forecasts. One critical concern with any forecast is the appropriate choice of the information set at any particular point in time; it would be inappropriate to estimate the model over the entire sample once and base forecasts on these estimates. This is because the Gibbs sampling algorithm that we use estimates the common factors backward through time and thus necessarily includes information from future time periods. To avoid this problem, we use a recursive forecasting procedure and estimate the econometric model first on a sample that ends in December 1929. With this sample, we produce forecasts of inflation rates for the following one, six, and twelve months. We subsequently add one month at a time to the original dataset and estimate the econometric model again. This approach ensures that we do not include information that was actually not available to agents into the information set. Furthermore, this approach allows the coefficients to be potentially different for each window. The model does not allow for structural breaks in the time series because structural break tests indicated breaks only in a very small number of inflation rates and growth rates of real variables.[10] Thus, it is reasonable to assume that a model with constant coefficients for each window is a satisfactory econometric model for

[10] Out of the 34 inflation time series available to us, the tests indicated structural breaks in only five. The only break close to the economic through was detected in clothing prices for households in July of 1932. Inflation in prices for agricultural machinery indicated a break in February of 1932. For the remaining three series - artificial fertilizers, building materials, and crafting materials - structural breaks were detected much earlier, in December of 1927, February of 1928, and July of 1927, respectively.

our question.

1.4.2 Data

We use a rich dataset from Wagemann (1935): 109 time series covering important areas such as production and employment, various price indices, trade, banking and monetary aggregates, and nominal interest rates. The data are collected at monthly frequency. All the time series, except for nominal interest rates, were seasonally adjusted prior to estimation and transformed into 100 times the monthly difference in natural logarithms of the adjusted series.[11] The nominal interest rate series are divided by 12 to convert them into monthly rather than annual interest rates. The panel with these definitions then starts in February 1925 and ends in June 1935. A complete list of the variables used to extract the common factors is provided in Appendix 1.B.3. When estimating the common factors, we used standardized values of the time series; thus, each series has zero mean and variance one for each estimation sample. This commonly used transformation of the data ensures comparability across time series. When constructing the forecasts, we convert them back by adding the mean and the standard deviation of each variable in every run of the sampling procedure to measure inflation expectations that can be compared with actually realized inflation rates.

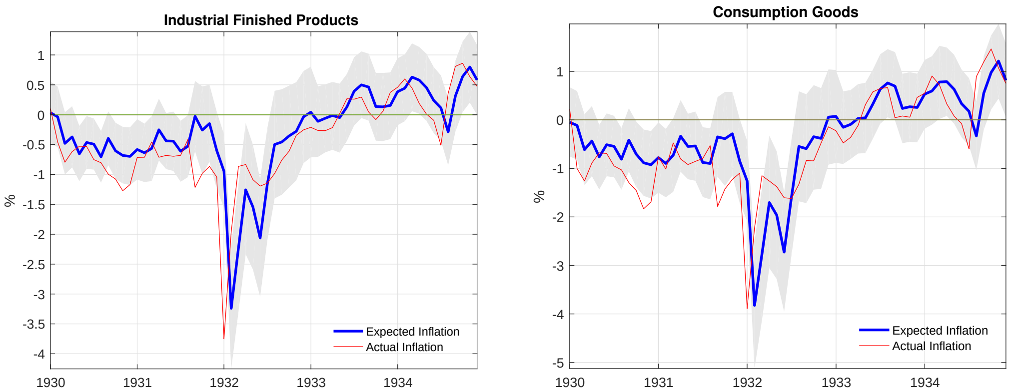
Notably, regarding data transformations, we convert the time series to month-to-month changes for two reasons. First, from the perspective of forming expectations about future inflation rates, it is much more sensible to assume that agents track the month-to-month change in price levels or aggregate activity, rather than assuming that the relevant growth rates span an entire year. Second, the underlying econometric theory is mostly developed for stationary time series, and most studies that forecast economic time series have used period-by-period log differences of the variables. Stock and Watson (1999), Stock and Watson (2002a), Stock and Watson (2002b), Eickmeier and Ziegler (2008) have all used transformations of this type, and we use this common practice in this paper.

1.4.3 Results

We start by considering one-month-ahead forecasts for two important inflation series, namely, inflation of industrial finished products and inflation of consumption goods. In this subsection, we report the median of the posterior distributions together with 95% of the posterior probability mass. Additionally, actual monthly

[11] Seasonal adjustments were performed with the Iris toolbox for Matlab, which implements X13-ARIMA routines for seasonal adjustments.

Figure 1.4: One-month ahead forecasts



Notes: Medians of posterior distribution shown in blue, 95% of the posterior probability mass in gray around the median. Realized inflation rates shown in red.

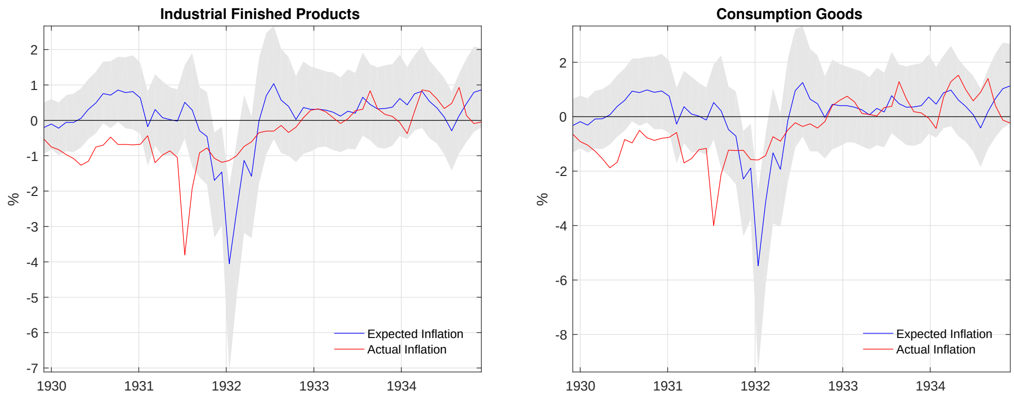
inflation rates are reported as a point of reference.[12] In Figure 1.4, the forecasts closely track the actual monthly inflation series, albeit with a one-month delay, and this is expected because of the linearity of the forecasting model. We observe two facts regarding expected inflation. First, throughout 1930 and 1931, continued deflation was expected each month, as shown by the consistent negative forecasted series. The large deflation shock of January 1932, which was due to the emergency decree in December 8, 1931, also fueled expectations of further large deflation rates. Given that by this time Germany had already experienced continued deflation rates pursued by the German administration even more strongly in December 1931, there is no reason for any difference in expectations.[13] Second, although the deflation period was to a large extent expected to continue during the early 1930s, the evidence for expected inflation rates after the summer of 1932 is mixed. Although inflation expectations together with actual inflation returned to lower deflation rates, both inflation and deflation are consistent with agents' forecasts from mid-1932 onward. Actual inflation rates were consistently negative until the end of 1932 for both series, and the evidence does not support the perspective that Germany experienced a similarly clear-cut reversal in inflation expectations as the U.S. did.

We next discuss forecasts over the following six months from each point in time in Figure 1.5. The two graphs show for each time point the expected inflation at six months from that particular point onwards. For example, the value for January 1931 measures expected inflation for July 1931 based on information available until

[12] Note that because we estimate the model on monthly inflation rates, these are of course smaller than annual inflation rates reported in Mitchell (1975). When we calculate annual inflation rates from our data, we find inflation rates that are very similar to those reported in Mitchell (1975). Figure 1.B.1 in Appendix 1.B.4 provides a comparison of these annual inflation rates as a consistency check.

[13] We provide a detailed description of the measures taken in December 1931 in Appendix 1.A.

Figure 1.5: Six-months ahead forecasts



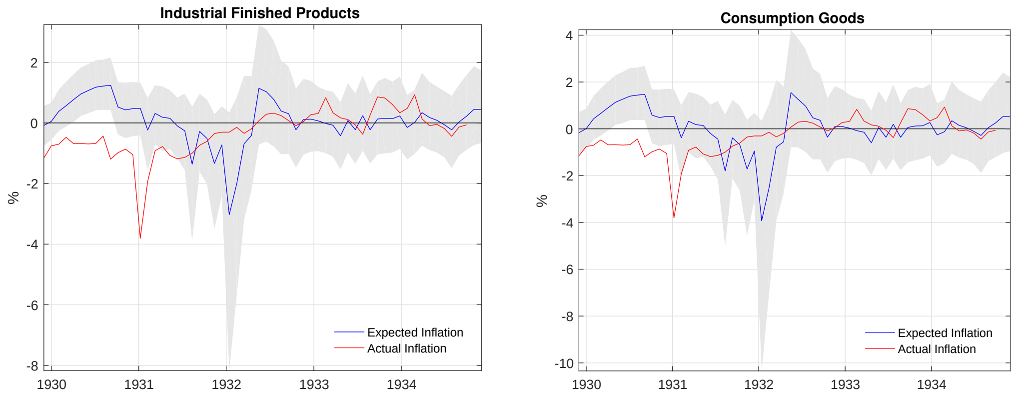
Notes: Medians of posterior distribution shown in blue, 95% of the posterior probability mass in gray around the median. Realized inflation rates shown in red.

January 1931. We also plot in red the realized inflation rates six months later. At first glance, the plot seems to indicate that over longer horizons, there was indeed an upward shift in inflation expectations starting in mid-1932; however, a comparison with expected inflation rates from the early 1930s shows that this is actually a reversal of expectations that the public already held, which was only disrupted by the abnormally large deflation period. This conclusion holds true for both inflation rates considered here, and both series are very similar.[14]

As a final check, we compute forecasts for an entire year from each time point onward; these results are shown in Figure 1.6. The pattern we observed for six months ahead expected inflation remains until mid-1932. The striking aspect about this figure, however, is the clear and quick convergence of expected inflation rates to zero toward the end of 1932. Essentially, there was nothing known to the public that would have indicated that positive inflation rates, if they were expected at all, would persist over time. Clearly, the results were opposite during the early years of the decade, where continued deflation impulses lead inflation expectations to be adjusted downward. This observation also suggests that inflation expectations are essentially zero toward the end of the sample period and to a large extent not driven by the stationarity assumption of the VAR model, which implies that over longer horizons, the forecasts converge to the unconditional mean of the factors. If this were the case, we would also see zero expected inflation at the beginning of the sample.

[14] We also estimated forecasts for other price series in the dataset and obtained similar results.

Figure 1.6: Twelve-months ahead forecasts



Notes: Medians of posterior distribution shown in blue, 95% of the posterior probability mass in gray around the median. Realized inflation rates shown in red.

1.4.4 Inflation expectations according to interest rate forecasts (Mishkin 1981)

Although a FAVAR model is a natural candidate from a forecasting standpoint, it is unlikely that economic agents at that time would have used such a complex model. Therefore, we corroborate our previous findings with simpler approaches. The first approach is from Mishkin (1981). His insight was that under the assumption of rational expectations, ex-ante real interest rates can be obtained as the fitted values from a regression of the ex-post real interest rate on a set of predictor variables:

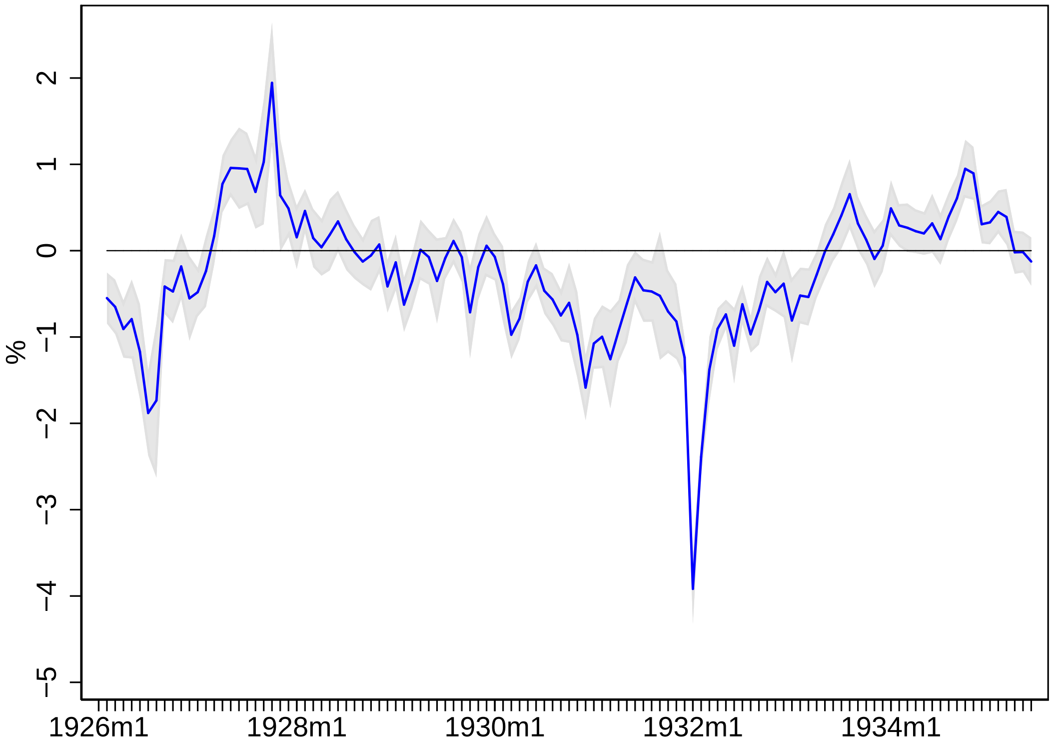
$$eprr_t = \beta X_{t-1} + u_t - \epsilon_t \tag{1.7}$$

With the nominal interest rate in hand, this allows to calculate expected inflation. This approach has also been used by Cecchetti (1992) and Romer (1992) for the U.S. Great Depression. In the spirit of Romer (1992), we run regressions of the realized monthly real interest rate on lagged values of the monthly growth rate of nominal total money in circulation, the monthly inflation rate of industrial finished products, the monthly growth rate of industrial production, and the Reichsbank nominal discount rate. For each variable, we include twelve lags in the regressions.[15]

Figure 1.7 shows the results of this exercise. Although inflation expectations were large and positive in 1927, deflation was expected from 1929 onward, in line with the results from the FAVAR approach. More importantly, for our present question, inflation expectations during 1932 were negative and very similar to those held by

[15] We use the following series, as stated in Wagemann (1935): nominal total money in circulation: XIII.b.1, “Geldumlauf insgesamt, Stand Monatsende”. Inflation rate of industrial finished products based on the price series: IX.B.23, “Industrielle Fertigwaren insgesamt”, Industrial production: III.B.11 “Industrieproduktion”, Reichsbank nominal discount: X.A.a.1, “Reichsbankdiskont”.

Figure 1.7: Expected inflation from real interest rate regressions



Notes: Implied expected inflation from regressions of monthly realized real interest rates on growth rates of nominal money in circulation and industrial production, inflation of industrial finished products, and the central bank nominal discount rate.

the public during 1930 and 1931. Only toward the end of 1933, we observe positive expected inflation rates. That is, the FAVAR and the real interest rate regressions deliver the same result, namely, the beginning of the German recovery cannot be explained by a dramatic shift in inflation expectations or positive expected inflation in general. Our results agree with Voth (1999) who found that during 1932, inflation expectations did not surpass their previous levels and only started to return to zero from 1933 onward.

1.4.5 Quantitative news estimates (Binder 2016)

Binder (2016) proposes to regress future inflation rates on a scaled measure of news counts. More specifically, the news measure is split into two components: one that counts news reports related to inflation ($News_t^+$), and one that measures the news count related to deflation ($News_t^-$). The log of the ratio is then used in the regression of future inflation on news. Formally, the model is

$$\pi_{t+j} = \alpha + \beta \ln \left(\frac{News_t^+}{News_t^-} \right) + u_t \tag{1.8}$$

The rationale underlying this approach is that news about price changes should be positively related to expected inflation rates. As proposed by Mishkin (1981), under rational expectations, it is possible to obtain estimates of expected inflation as the fitted values of future inflation rates on the news variable. Table 1.2 shows the results of this approach where we used one month (columns 1 to 2) and 12 months (columns 4 to 5) ahead inflation rates for industrial finished products and consumption goods as dependent variables. As expected, the news variable is highly significantly correlated with future inflation rates with a positive coefficient. Notably, when we replace future inflation rates with future growth in industrial production (columns 3 and 6), news about price changes are not significant and explain very little of the variation.

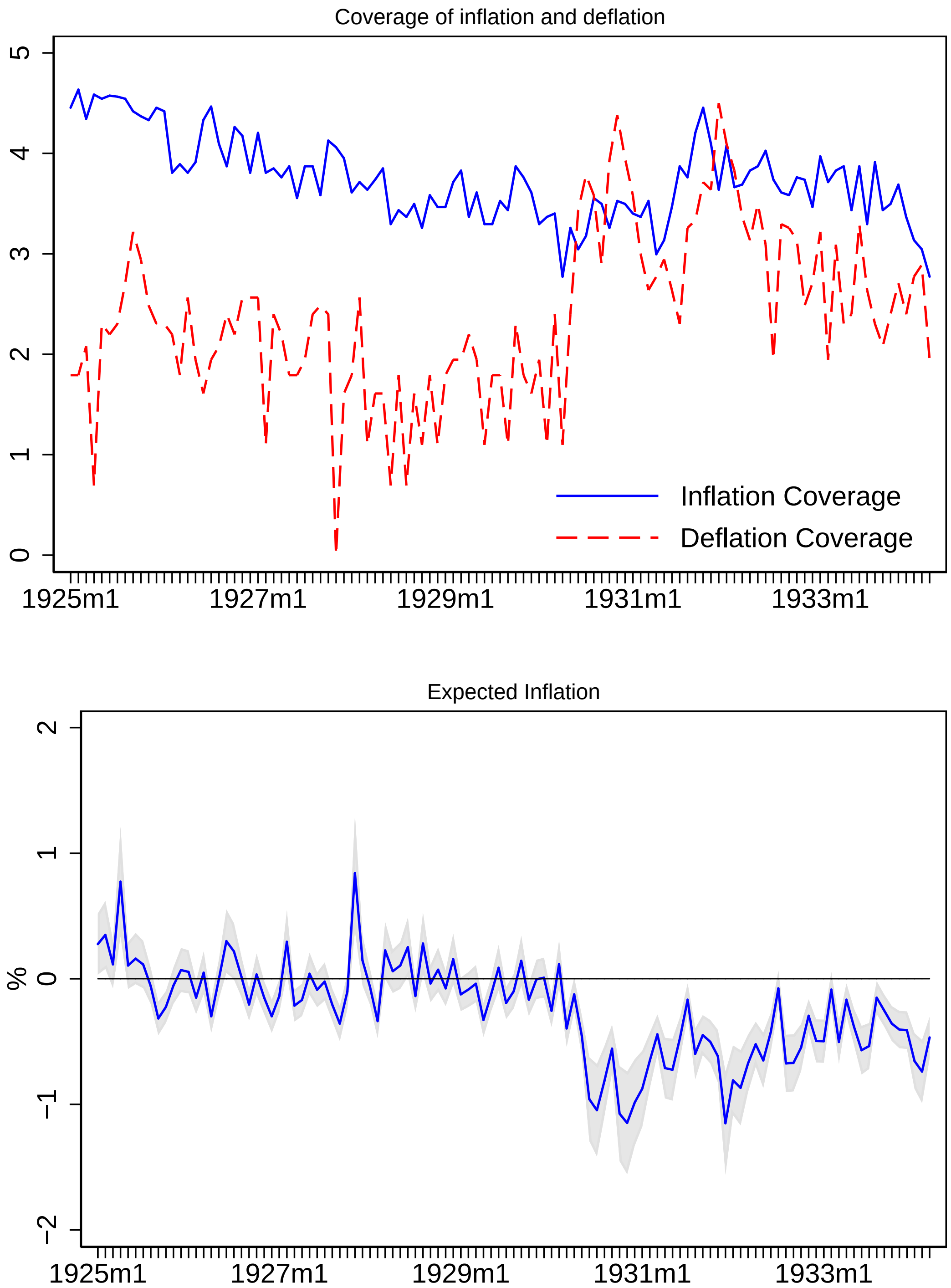
Table 1.2: Forecasting regressions with narrative news measure

	π_{t+1}^{IFP}	π_{t+1}^{C}	ΔIP_{t+1}	π_{t+12}^{IFP}	π_{t+12}^{C}	ΔIP_{t+12}
News	0.405***	0.436***	0.321	3.397***	3.958***	3.336
	(0.0864)	(0.0996)	(0.367)	(0.437)	(0.557)	(1.984)
Constant	-0.802***	-0.933***	-0.511	-8.764***	-10.77***	-7.579*
	(0.140)	(0.163)	(0.633)	(0.839)	(1.075)	(3.278)
N	110	110	110	99	99	99
R^2	0.238	0.178	0.00650	0.217	0.184	0.0251

Notes: Regressions of future inflation rates and growth of industrial production on the news measure constructed under Equation 1.8. π_{t+j}^{IFP} and π_{t+j}^{C} denote inflation rates of industrial finished products and consumption goods, respectively, between periods t and $t+j$. ΔIP_{t+j} denotes the growth rate in industrial production between periods t and $t+j$. Series of prices and industrial production are taken from Wagemann (1935), series No. III.B.11, IX.B.23, IX.B.30.

Further, in Figure 1.8, we plot the two components of the news measure in logs and expected inflation of industrial finished products as measured by the fitted values of column 1 of Table 1.2. As shown in the upper part, the newspaper coverage of deflation increases dramatically during the final months of 1930, thus closing the previous gap between the number of articles covering inflation and deflation. From this time onward, many articles addressed the prevalent and ongoing price reductions of many products. The deflation news count then remains elevated throughout, similar to the inflation news count. It is crucial to note here that deflation is not mentioned in the context of deflation coming to an end, but rather is described as

Figure 1.8: News coverage and expected inflation



Notes: The first panel plots the log of the number of articles covering inflation and deflation. The second panel plots the fitted values from the first column of Table 1.2.

a continued process of price cuts. That is, the deflation news count does not serve as a double negative for inflation. In summary, the regression results imply that expected inflation was actually negative from mid-1930 onward, with no sign of a shift toward smaller deflation rates or even inflation rates.

The use of only one newspaper in this context is potentially problematic. Because different newspapers favor different political positions, how the same events are presented and discussed across newspapers may differ dramatically. Hence, we conducted inflation and deflation news counts for four additional newspapers and constructed a scaled news measure for each them. A problematic issue is that sometimes the additional newspapers were not available or started at later dates. These measures are shown in Appendix 1.C.4, Figure 1.C.4. The scaled measure is remarkably similar across newspapers, and the results from model (1.8) are robust to this broader news measure. This result suggests that our focus on only one newspaper does not bias our results, and that our baseline news count accurately captures the media account regarding inflation and deflation at that time.

From the results of this section, across approaches, we consistently observe evidence that during 1932, expected inflation rates were stable and negative. That is, when the recovery started, there is little to suggest that inflation expectations changed. We interpret the large decline in expected inflation rates in early 1932 following the price cuts dictated by the government as a very temporary negative shock that did not affect inflation expectations much during the following months. These results are also consistent with our narrative analysis. Newspaper articles reveal strong opposition from policymakers to inflationary policies, and a sharp increase in newspaper coverage of deflation from 1931 onward. Apart from the sharp fall of prices in January 1932 because of forced price reductions due to the emergency decree, the time series forecasts indicate essentially the same expected price movements in 1932 as in 1931: no shift to inflation or flat expectations.[16]

[16] We further considered additional approaches to identify inflation expectations, namely using the spot and forward exchange rates (as proposed by Voth (1999)) or inferring information regarding expected inflation from shipment rates on railroads that Klug et al. (2005) tested for the U.S. interwar period. However, because the data were not available, we could not pursue both approaches. For instance, the Reichsmark forward exchange data ended in August 1931. Railroad freight rates and shipper's forecasts for the German interwar period are yet to be discovered. Lastly, referring to Hamilton (1992), we investigated the use of futures prices for wheat to infer a market-based measure of expected inflation rates. For this purpose, from the *Vossische Zeitung*, we collected monthly wheat futures of different maturity in the months they were posted between 1925 and 1933. An issue about this data is that it is very frequently missing; thus, we were unable to construct a continuous series. In addition, a potential caveat is that commodity futures prices potentially reflected exchange rate fluctuations or global price movements. We therefore decided to exclude them from this discussion; however, we added them as an additional section of Appendix 1.B.5, where we plot the prices and expectations for each available month and duration in Figures 1.B.2 and 1.B.3, respectively.

1.5 Role of inflation expectations for the recovery

Thus far, our results imply that compared with the U.S., no measures to increase expected inflation were feasible in Germany. The quantitative forecasting results show no indication for a clear shift in inflation expectations at any point between 1931 and 1933. The detailed narrative study of media articles ruled out a regime change that was potentially undetected by our quantitative estimates. Newspaper article counts with respect to inflation reveal four spikes in the coverage, which could possibly indicate inflationary news.

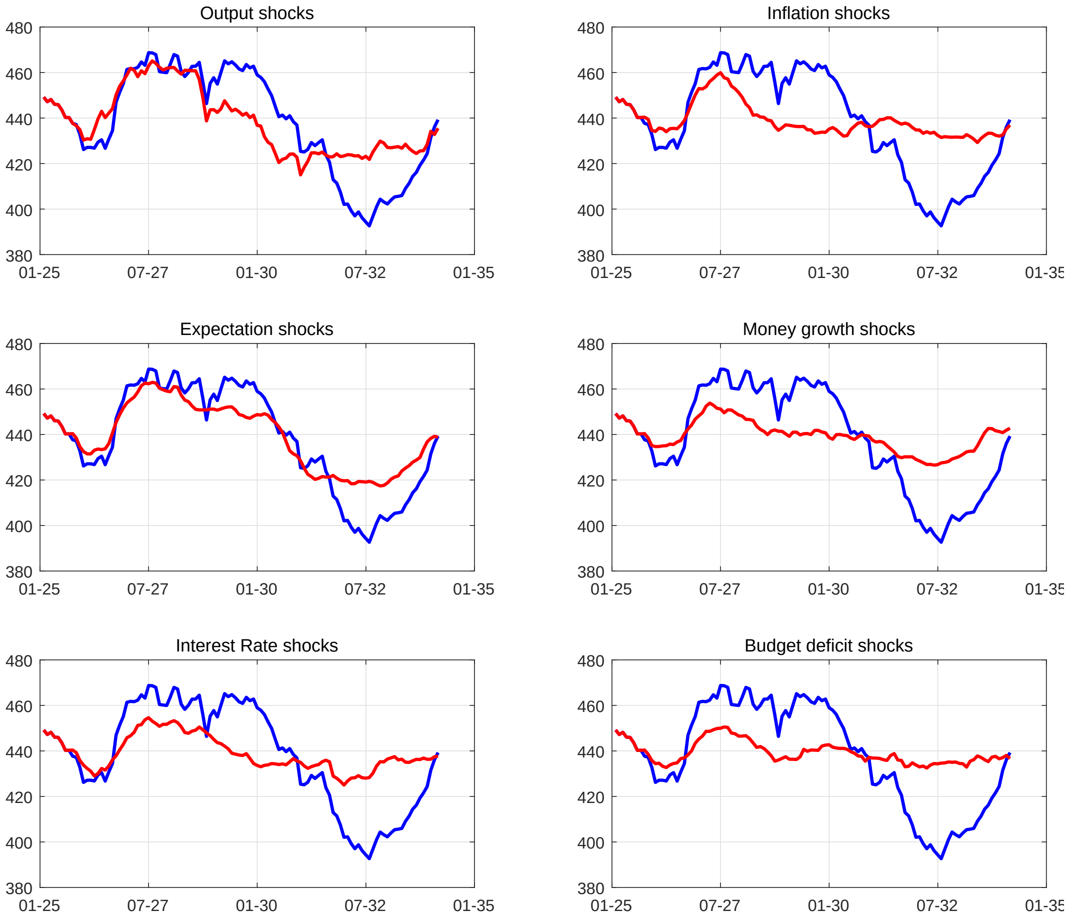
To provide further evidence for the absence of an inflation expectations channel for the recovery, we estimated a small-scale VAR model of the form

$$Y_t = C + \sum_{l=1}^{L} B_l Y_{t-l} + u_t \tag{1.9}$$

with the following variables in order: the log of industrial production, the log of the price index of industrial finished products, expected inflation, the nominal central bank interest rate, the growth rate of nominal money balances, and the real government budget deficit. As our measure of expected inflation, we use the expected inflation series as implied by the approach in Binder (2016). The government deficit is obtained from Wagemann (1935) as the difference between total expenditures and revenues for the central government. We posit that this specification captures several potential channels through which the recovery may have occurred. The VAR setting allows us to assess the contribution of structural shocks to each variable on the path of the real economy, in this case, industrial production, through historical decomposition. This approach was similarly employed by Shibamoto and Shizume (2014) to study the role of inflation expectations for the case of Japan. The structural shocks are identified through a Cholesky factorization of the residual covariance matrix with the ordering just specified. The deficit is ordered last because total expenditures naturally include automatic stabilizers that react contemporaneously to the real economy, and because fiscal policy is assumed to react to the stance of monetary policy. We also assume that shocks to policy variables do not affect the real economy and expectations within each month but that policy variables react contemporaneously to shocks to the real economy. This ordering follows Shibamoto and Shizume (2014) who also order expected inflation after real economy variables but before policy variables in their VAR analysis.[17] The lag length was chosen to be

[17] The results are robust to ordering expected inflation first or last in the VAR. Additionally, the results are virtually identical when we use expected inflation as implied by real interest rate regressions as in Romer (1992) or Mishkin (1981). Robustness checks can be found in Appendix 1.B.6.

Figure 1.9: Historical decomposition of industrial production



Notes: Blue lines denote 100 times the log of the industrial production index. Red lines depict the series for industrial production that would have been obtained if only shocks to the respective variable occurred. Sources: Wagemann (1935), series No. III.B.11, IX.B.23, X.A.a.1, XIII.b.5, XVIII.A.b.5, XVIII.A.b.6.

eight to strike a balance between a large number of several estimation parameters with a short sample and to accurately capture the dynamics of the variables over time.

The result is presented in Figure 1.9. In each panel, we plot the original series for industrial production in blue, and the time series that we obtained had only structural shocks to the respective variable occurred over time in red. We observe that together with shocks to the real economy, shocks to expected inflation were important drivers of the economic downturn until mid-1931. This result is consistent with the perspective that the public, expecting deflation throughout this period, postponed consumption, leading to the large economic decline. The second large drop during the banking crisis is largely attributed to shocks to the nominal rate and, to a lesser extent, to fiscal policy. This result agrees with the arguments in the literature that the banking crisis forced the German monetary authority to drastically increase the interest rate to prevent capital outflows under the gold standard. Additionally, the fiscal authority continued the highly unpopular austerity policies, which have often been blamed for worsening the Depression. Lastly, we observe that the start of the recovery emanated from the nominal and real sector, whereas expected inflation became an important economic driver only later on, once the recovery had been under way for some months. This result is consistent with our previous findings that expected inflation was stable and negative during 1932.

1.6 Implications and future directions

Our results rule out a shift in inflation expectations as a potential explanation of the German recovery from the Great Depression. The narrative account and time series models show that when the recovery started, expected inflation rates remained unchanged. Whenever fears of inflation were mentioned, politicians denounced any price-increasing policies and emphasized the unconditional stability of the currency. Therefore, the exact cause of the recovery remains a puzzle. Our results from the historical decomposition support some of the explanations that have been previously analyzed. One potential explanation is that the conference of Lausanne boosted public confidence or significantly reduced uncertainty about future economic policy because it put an end to the issue of reparation payments that plagued German politics. Based on an indicator of the state of business across major German industries, Buchheim (2003, 2008) showed that producer confidence indeed markedly increased during the second half of 1932. Machine orders lastingly increased and stock markets had been uptrending since June. Voth (1999) made the same argument when he found that fears of inflation during the Great Depression greatly limited the scope for expansionary policies, at least before the resolution of the reparations problem. Tooze (2006) similarly views the agreement reached in Lausanne as a likely precondition for the recovery because it provided the financial relief that banks and the government needed to promote active economic programs. The narrative sources, however, contain no evidence of a link to an increase in prices.

Temin (1990) provided a monetary explanation for the successful German recovery and argued that low wages played a crucial role in the German economic success of the 1930s. The freezing of wages allowed for a quick reutilization of underused resources. In contrast to raising demand and thereby prices of stimulated production, political pressure under the Nazis reallocated production to the military and public sector. The now controlled economy no longer depended on the price mechanism for allocation of scarce resources. In that sense, the Nazis managed potential inflationary tendencies by freezing prices. As Temin (1990) argued, instead of raising prices after 1933, producers started to reduce the quality of goods, which underlines our argument, that is, inflation expectations likely played no central role at the beginning of the German recovery.

In many industries, prices were fixed by cartels and government controls, so we surveyed our narrative sources for signs of expected rationing as a potential explanation of the recovery, with no indications that this channel was at play. In contrast, in 1932 and 1933, we find numerous pieces of evidence indicating unused capacities on a large scale that could have quickly been put into operation in case of surges in demand. In addition, during the Great Depression, the data from Buchheim (2008) indicate that low consumer demand led to growing inventories and to businesses reducing employment. Hence, we deem it unlikely that households expected shortages at the beginning of the recovery, causing an increase in demand and production. Although rationing certainly became a problem in the roaring economic upturn during the mid-1930s, it did not play a role in jump-starting the recovery.

By contrast, fiscal policy serves as a promising alternative explanation of the recovery period. In August of 1932, the newly elected chancellor Franz von Papen publicly announced his fiscal program. The main feature of the program was tax rebates and wage cuts, and both were intended to incentivize businesses to hire workers. Papen declared that deflation, if not stopped, would further harm the economy. Although this announcement could have changed public expectations about future inflation and stimulated the economy this way, the program did not signal the intention to raise prices or reduce price reductions. This stands in stark contrast to Roosevelt's pledges to raise prices. For the U.S., the economic reform foresaw the devaluation of the U.S. dollar, and unprecedented discretionary powers of U.S. monetary policy. In Germany, by contrast, the chancellor explicitly stated that the Reichsmark would not be devaluated, and his program foresaw no changes regarding the regulations of the German central bank, which had hindered it from acting as a financier of previous public works projects. The central bank only gave the promise to grant limited credit to the government to finance tax redemptions. The repayment of the Reichsbank loan was described well in the program and in the news accounts, indicating that the potential monetary expansion was restricted to a limited time period. In summary, the Papen program did not work through affecting inflation expectations as the monetary expansion in the U.S. did. Instead it attempted to establish incentives for the private sector, even at the cost of further wage cuts argued to be indispensable. Examining the real effects of the Papen program in detail is a promising topic to explore in further research.

Jalil and Rua (2016) thoroughly traced the causes of the regime shift in the U.S. to inflation expectations in spring 1933, and our study provides considerable evidence that no such event occurred in Germany. The sources of the German recovery remain unclear. The story of Temin and Wigmore (1990) and Eggertsson (2008) may hold for the U.S., but in the German recovery, inflation expectations played no major role.

APPENDICES

1.A Price cuts under the emergency decree of December 1931

The source for the drastic decrease in prices in January 1932 indicated by our time series evidence was because of one policy measure: the concerted and official reductions in prices, rents, wages, and interest rates based on the presidential decree of December 8, 1931. Because the timing and the expectations of the measure are important for our reasoning, we provide a detailed description that relies on news articles and the official legislative texts in the law gazette of the Reich ("Reichsgesetzblatt" = RGB). The decree ("Vierte Verordnung des Reichspräsidenten zur Sicherung von Wirtschaft und Finanzen und zum Schutze des inneren Friedens") stated that fixed prices had to be reduced by at least 10% by January 1, 1932, compared with the prices on June 30, 1931. As fixed prices, the law defined prices of products established by contractual agreements such as cartels or syndicates. The industries for which this was the case were iron and metal, construction, chemicals, paper, glassware, ceramics and textiles, and coal (RGB 1931, Nr. 79, 1. Teil, Kap. 1.1 §1(2) and §3). The reductions were complemented by a commissioner of price control for which a second decree was issued on December 8. The commissioner's task was to enforce the consumer price reductions and punish violations (see "Verordnung über die Befugnisse des Reichskommissars für Preisüberwachung", RGB 1931, Nr. 80). The price reductions were accompanied by reductions in rents by 10% and wages by up to 15%. The presidential decree also contained a section on interest rates. The nominal rates on many bonds, obligations, and debt instruments were reduced by a proportionate share between zero and up to 50% (see RGB 1931, Nr. 79, parts 1-3). To support the intended effects of the presidential decrees, the Reichsbank reduced the discount and Lombard rates on December 10, 1931. The news media covered the presidential decree and its consequences in detail after the decree was published. Before its passage, the exact measures were unknown. The first mention of possible future price and wage cuts was on December 3. Nonetheless, on December 8, *Vossische Zeitung* reported of "an emergency decree with surprises", because the decree had become the subject of last-minute changes. After December 8, the measures were discussed in-depth and should thus be regarded as common knowledge in December 1932 and to have taken effect on January 1, 1932. The first price reductions in response to the decree were mentioned on December 18 (see *Vossiche Zeitung*, *Berliner Lokal-Anzeiger* and other newspapers, November 1931 through January 1932).

1.B Time series evidence supplements

1.B.1 Prior distribution

We specify the following prior for the parameters in each equation $i = 1, \ldots, N$ of (1):

$$\lambda_i \sim \mathcal{N}(\underline{M}, \underline{V}) \tag{1.10}$$

$$\underline{M} = 0_{K \times 1} \tag{1.11}$$

$$\underline{V} = I_K \tag{1.12}$$

Next, the idiosyncratic variances follow

$$\sigma_i^2 \sim \mathcal{IG}\left(\frac{\alpha}{2}, \frac{\delta}{2}\right) \tag{1.13}$$

$$\alpha = 6 \tag{1.14}$$

$$\delta = 0.001 \tag{1.15}$$

These values are from Ritschl and Sarferaz (2014). For the VAR block of the model, we set the natural conjugate Normal-Inverse-Wishart prior as in Kadiyala and Karlsson (1997)

$$\Sigma \sim \mathcal{IW}\left(\underline{S}, K + 2\right) \tag{1.16}$$

$$vec(B)|\Sigma \sim \mathcal{N}(vec(\underline{B}), \Sigma \otimes \underline{G}) \tag{1.17}$$

The matrix $\underline{S}$ is the K-dimensional identity matrix. The diagonal elements of $\underline{G}$ are chosen such that the prior variance on the parameter of variable j in equation k at lag l is $p\frac{\sigma_k^2}{l\sigma_j^2}$, where $p = 0.05$. The prior covariance matrix of the VAR parameters is given by $\underline{S} \otimes \underline{G}$. The posterior distributions are shown in Appendix 1.B.2. $\underline{B}$ is the $KL \times K$ matrix of prior VAR parameters, which has all zeros.

1.B.2 Estimation of the FAVAR model

We perform an estimation using the following steps. We start with a sample that ends in December of 1929 and estimate the factor model on standardized data up to that point in time. After an initial burn-in period, we estimate our forecasts for one, six, and twelve months according to Equations (1.5) and (1.6). Next, we expand the sample by one month and perform the sampler and regressions again. This process is performed until December 1934. In principle, an estimation of the factor model over the entire sample once is possible, but this creates the problem that when we sample the common factors, the algorithm moves backward through the sample and draws the factors based on their conditional distributions. Thus, for each time period we use information that economic agents would not have known yet. Our procedure, instead, uses only available information when estimating the model.

For each run, we perform 50,000 iterations of the sampler, and each time, we reset the prior covariances in the VAR to account for the changing information available. Next, we discard the initial 30,000 draws and base the inference on the remaining draws. Our point estimates are then the medians of the posterior distributions at each point in time together with 95% bands of the posterior probability mass. We use $L = 12$ in the estimation and $K = 3$ common factors for each run.
We start by drawing the idiosyncratic variances conditional on the data, factors, and loadings from the following distribution

$$\sigma_i^2 \sim \mathcal{IG}\left(\frac{T+\alpha}{2}, \frac{\epsilon_i'\epsilon_i + \delta}{2}\right) \tag{1.18}$$

where ϵ_i is the $T \times 1$ vector of residuals in equation i. Next, based on the drawn variances we sample the factor loadings from

$$\lambda_i \sim \mathcal{N}(\overline{M}, \overline{V}) \tag{1.19}$$

$$\overline{V} = \left(\underline{V}^{-1} + \frac{1}{\sigma_i^2} f'f\right)^{-1} \tag{1.20}$$

$$\overline{M} = \overline{V}\left(\underline{V}^{-1}\underline{M} + \frac{1}{\sigma_i^2} f'Y_i\right) \tag{1.21}$$

where f is $T \times K$ and Y_i is $T \times 1$. To sample the VAR block, we write the VAR as

$$F = XB + U \tag{1.22}$$

where F is the $(T - L) \times K$ matrix of common factors; X is the $(T - L) \times KL$ matrix of the lagged common factors that has the first lags of all factors in the first K columns, the second lags in the next K columns, and so forth, and B is the $KL \times K$ matrix of coefficients. We then draw the VAR covariance matrix and the coefficients from

$$\Sigma \sim \mathcal{IW}(\overline{S}, T - L + K + 2) \tag{1.23}$$

$$vec(B) \sim \mathcal{N}(vec(\overline{B}), \Sigma \otimes \overline{G}) \tag{1.24}$$

where the matrices in these distributions are defined as

$$\overline{G} = \left(\underline{G}^{-1} + X'X\right)^{-1} \tag{1.25}$$

$$\overline{B} = \overline{G}\left(\underline{G}^{-1}\underline{B} + X'X\hat{B}\right) \tag{1.26}$$

$$\overline{S} = \underline{S} + \hat{B}'X'X\hat{B} + \underline{B}'\underline{G}^{-1}\underline{B} + \hat{U}'\hat{U} - \overline{B}'\left(\underline{G}^{-1} + X'X\right)\overline{B} \tag{1.27}$$

and variables with a hat denote ordinary least squares' quantities.
Sampling the latent factors is based on the state-space representation of the model; we write this as

$$Y_t = C + HF_t + \epsilon_t \tag{1.28}$$

$$F_t = \tilde{B}F_{t-1} + \tilde{v}_t \tag{1.29}$$

$$H = \begin{bmatrix}\Lambda & 0_{N\times K(L-1)}\end{bmatrix} \tag{1.30}$$

$\tilde{B}$ denotes the companion form matrix of the VAR, and $\tilde{v}_t$ are the original errors in the VAR, appended with zeros. To sample the common factors, we employ the Carter-Kohn algorithm. We first use the Kalman filter to produce estimates of the state vectors and then go backward in the sample; at each point, we draw the common factors from the corresponding normal distribution. The exact procedure is explained in Kim and Nelson (1999).

1.B.3 Variables to estimate the common factors

The following table lists all variables used to extract the common factors and transformations performed prior to estimation.

Table 1.B.1: Data and factors table

Variable	Code	SA	Transformation
Production and Employment			
Number of Uneymployed	II.B.20	1	1
Number of Recipients of Unemployment Benefits	II.B.21	1	1
Industrial Production	III.B.11	1	1
Production goods	III.B.12	1	1
Investment goods	III.B.13	1	1
Consumption goods, elastic demand	III.B.16	1	1
Consumption goods, inelastic demand	III.B.17	1	1
Fuels	III.B.18	1	1
Basic materials	III.B.19	1	1
Constructions	III.B.20	1	1
Iron	III.B.22	1	1
Construction Industry total	III.B.26	1	1
Coal	III.B.31	1	1
Gas	III.B.32	1	1
Electricity	III.B.33	1	1
Oil	III.B.35	1	1
Paper	III.B.36	1	1
Potassium	III.B.37	1	1
Textiles	III.B.39	1	1
Shoes	III.B.40	1	1
Household Porcelain	III.B.42	1	1
Orders of machinery within Germany	IV.4	1	1
Imports			
Foods and Drinks	VII.B.a.5	1	2
Raw materials and semi-finished products	VII.B.a.6	1	2
Finished products	VII.B.a.7	1	2
Exports			
Foods and Drinks	VII.C.a.13	1	2
Raw materials and semi-finished products	VII.C.a.14	1	2
Finished products	VII.C.a.15	1	2

Notes: "Code" refers to the numeric codes from Wagemann (1935). "SA" denotes whether series was seasonally adjusted (1) or not (0). "Transformation" denotes whether a series entered in levels (0), transformed into 100 times first differences of logs (1), or converted into real terms using the cost of living index and then transformed into 100 times first differences of logs (2).

Table 1.B.1 Continued

Variable	Code	SA	Transformation
Prices			
Agricultural products	IX.B.3	1	1
Herbal foods	IX.B.4	1	1
Animals for slaughter	IX.B.5	1	1
Cattel products	IX.B.6	1	1
Animal feed	IX.B.7	1	1
Colonial products	IX.B.8	1	1
Industrial raw materials and semi-finished products	IX.B.9	1	1
Coal	IX.B.10	1	1
Iron	IX.B.11	1	1
Metal	IX.B.12	1	1
Textiles	IX.B.13	1	1
Leather	IX.B.14	1	1
Chemicals	IX.B.15	1	1
Artifical fertilizers	IX.B.16	1	1
Power oils and lubricants	IX.B.17	1	1
Rubber	IX.B.18	1	1
Paper	IX.B.19	1	1
Building materials	IX.B.20	1	1
Basic materials	IX.B.21	1	1
Industrial raw materials for exports	IX.B.22	1	1
Industrial finished goods	IX.B.23	1	1
Means of production	IX.B.24	1	1
Agricultural dead inventory	IX.B.25	1	1
Agricultural machinery	IX.B.26	1	1
Commercial facilities	IX.B.27	1	1
Commercial machinery	IX.B.28	1	1
Crafting materials	IX.B.29	1	1
Consumption goods	IX.B.30	1	1
Furniture	IX.B.31	1	1
Clothing and Shoes	IX.B.33	1	1
Nutrition	IX.C.36	1	1
Heating and Lighting	IX.C.38	1	1
Clothing for households	IX.C.39	1	1
Other	IX.C.40	1	1
Nominal Interest Rates			
Reichsbank discount rate	X.A.a.1	0	0
Reichsbank lombard rate	X.A.a.2	0	0
Credit costs	X.A.b.5	0	0
Daily due money	X.A.b.8	0	0
Daily money	X.A.c.14	0	0
Monthly deposits	X.A.c.15	0	0
Private discount rate	X.A.c.16	0	0
Goods of Exchange	X.A.c.17	0	0

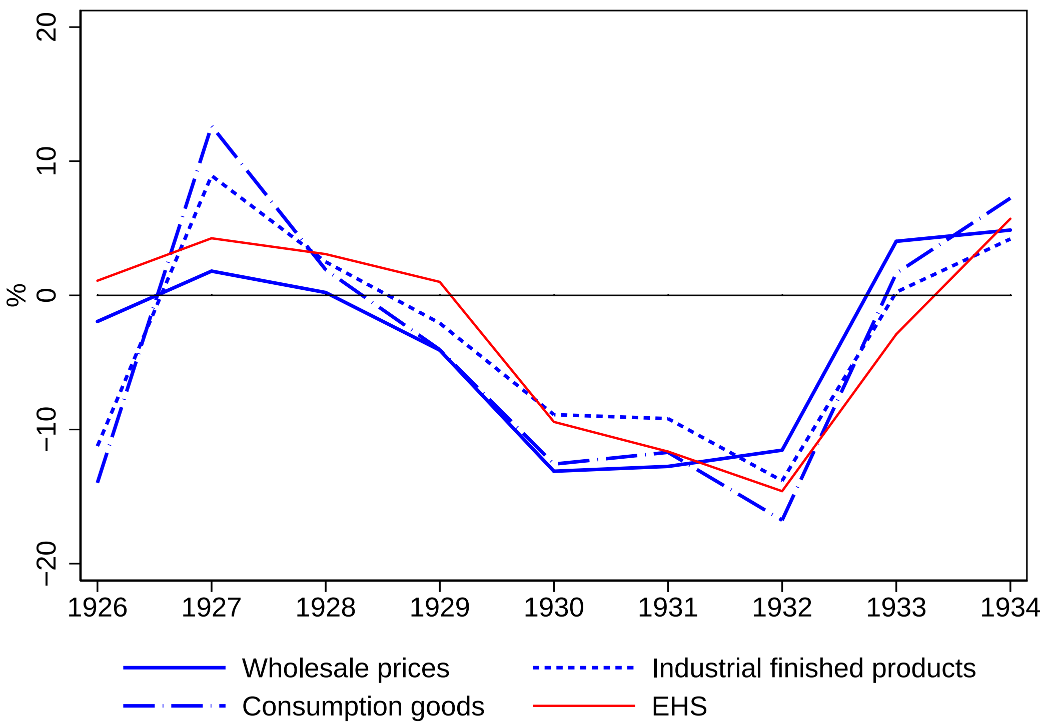
Table 1.B.1 Continued

Variable	Code	SA	Transformation
Sales			
Foods	VI.A.2	1	1
Textiles and Clothing	VI.A.3	1	1
Furniture and Household Appliances	VI.A.4	1	1
Grocery stores	VI.A.7	1	1
Rural grocery stores	VI.A.8	1	1
Drugstores	VI.A.12	1	1
Men clothing	VI.A.14	1	1
Women clothing	VI.A.15	1	1
Shoes	VI.A.16	1	1
Furniture specialist shops	VI.A.19	1	1
Colonial products, central retailer cooperative	VI.B.25	1	1
Colonial products, local retailer cooperative	VI.B.26	1	1
Drugs, retailer cooperative	VI.B.27	1	1
Furniture, retailer cooperative	VI.B.28	1	1
Clocks, retailer cooperative	VI.B.29	1	1
Innkeeping, retailer cooperative	VI.B.30	1	1
Bakeries, central retailer cooperative	VI.B.31	1	1
Bakeries, local retailer cooperative	VI.B.32	1	1
Butchers, retailer cooperative	VI.B.33	1	1
Cutters, retailer cooperative	VI.B.34	1	1
Shoemakers, retailer cooperative	VI.B.35	1	1
Saddlers, retailer cooperative	VI.B.36	1	1
Painter, retailer cooperative	VI.B.37	1	1
Woodworking, retailer cooperative	VI.B.38	1	1
Metalworking, retailer cooperative	VI.B.39	1	1
Money and Banking			
Circulation of money	XIII.b.5	1	2
Circulation of notes (Reichsbank and private)	XIII.b.6	1	2
Circulation of Rentenbank notes	XIII.b.7	1	2
Circulation of coins	XIII.b.8	1	2
Billing traffic by the Reichsbank	XIII.c.9	1	2
Giro traffic by the Reichsbank	XIII.c.10	1	2
Post check traffic	XIII.c.12	1	2
Credit on post check accounts	XIII.c.13	1	2
Drawings of bills	XIII.c.16	1	2
Circulation of bills of exchange	XIV.A.3	1	2
Stock of gold and currency, Reichsbank	XIV.B.5	1	2
Investments, Reichsbank	XIV.B.6	1	2
Exchange loans, Reichsbank	XIV.B.7	1	2
Deposits, Reichsbank	XIV.B.8	1	2

1.B.4 Price indices in comparison

Figure 1.B.1 shows, as a consistency check, a comparison between our inflation data and wholesale price inflation, as stated in Mitchell's (1975) European historical statistics (EHS). Because the data in Mitchell (1975) are annual, we calculate implied annual inflation rates from our monthly series. Clearly, the data are consistent across both sources.

Figure 1.B.1: Comparison of annual inflation rates

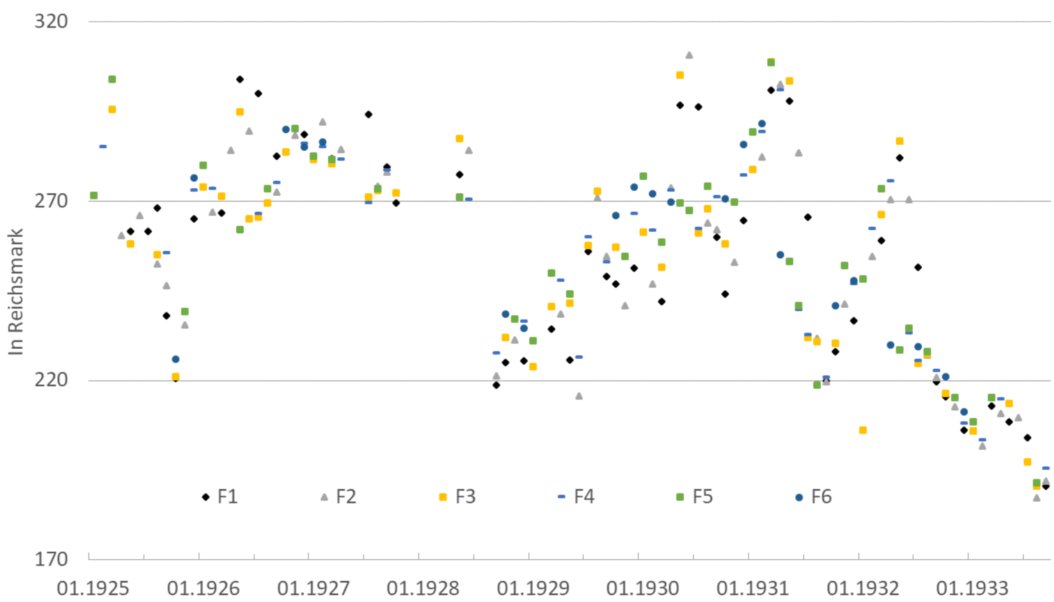


Notes: Blue lines depict annual inflation rates implied by the monthly data used in the main text. As a comparison, the red line depicts annual wholesale price inflation as stated in the European historical statistics (Mitchell, 1975).

1.B.5 Futures data

As mentioned in Subsection 1.4.5 of the main text, we also considered futures prices as indicator for inflation expectations. From the *Vossische Zeitung*, we collected monthly wheat futures of different maturity in the months they were available between 1925 and 1933. We plot the prices for each available month and duration in Figure 1.B.2. A problematic issue is that some of the data are frequently missing; thus, we could not construct a continuous series. In addition, a potential caveat is that commodity futures prices potentially reflected global price movements rather than changes in domestic expectations (see Albers (2018) and O'Rourke (2000) on the interwar commodity market integration). As such, this data is unlikely to be informative regarding the question addressed in this paper.[18]

Figure 1.B.2: Wheat futures prices



Notes: Futures prices on wheat 1925 to 1933 in Reichsmark per metric ton. F1 to F6 denote the futures prices at the beginning of each month for the contracts with settlement dates 1 to 6 months in the future. For months in which they were traded, each point denotes the first future price of the month mentioned in the business section of *Vossische Zeitung*. For example, F1 denotes the price of a contract settled at the end of the month, e.g., a futures price F1 of September 1 denotes a contract to be settled by September 30, and a price F2 on September 1 denotes a contract to be settled by the end of October of the same year. No futures prices were reported after September 1933.

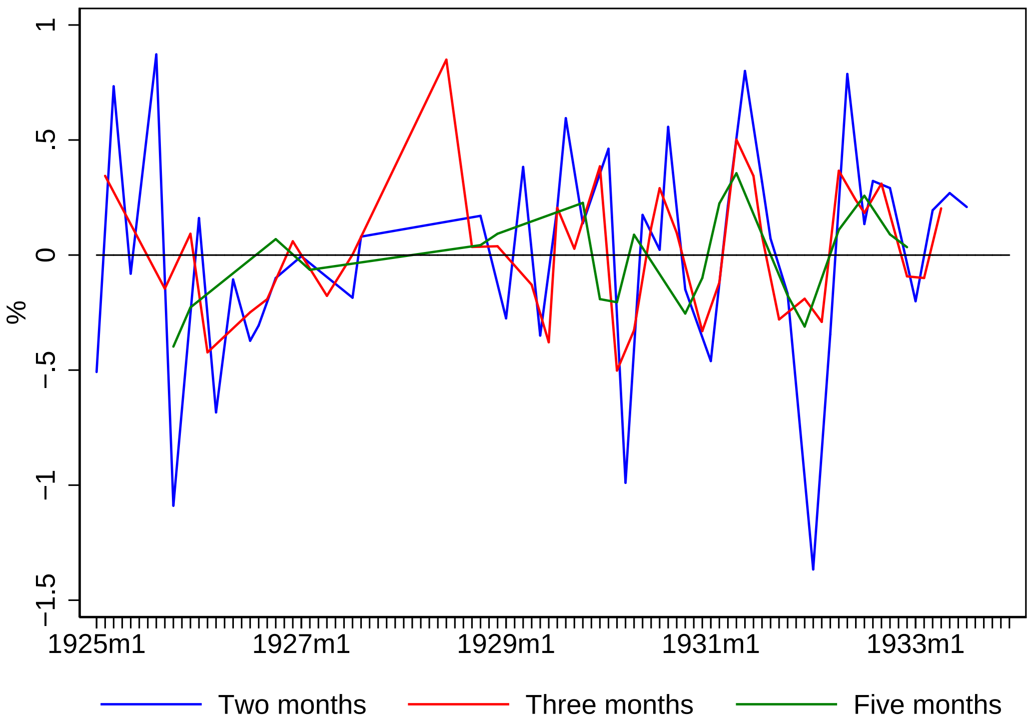
According to Hamilton (1992), these data could enhance the understanding of inflation because the difference between futures prices and realized spot prices should, under rationality, reveal traders' expectations of future inflation, that is,

[18] We thank the referees for pointing out this caveat.

$$\pi^e_{j,t} = \frac{12}{T}\left(\log(F_{j,t}) - \log(S_{j,t+s})\right) \tag{1.31}$$

where $F_{j,t}$ is the future price of commodity j at time t, to be delivered at time $t+s$. $S_{j,t+s}$ is the spot price of the same commodity j. As noted in the literature, instead of the spot price, we use future prices with differing durations to settlement date to remove transaction costs from the calculation (see also Voth (1999) for further details).

Figure 1.B.3: Implied inflation expectations from futures prices on wheat



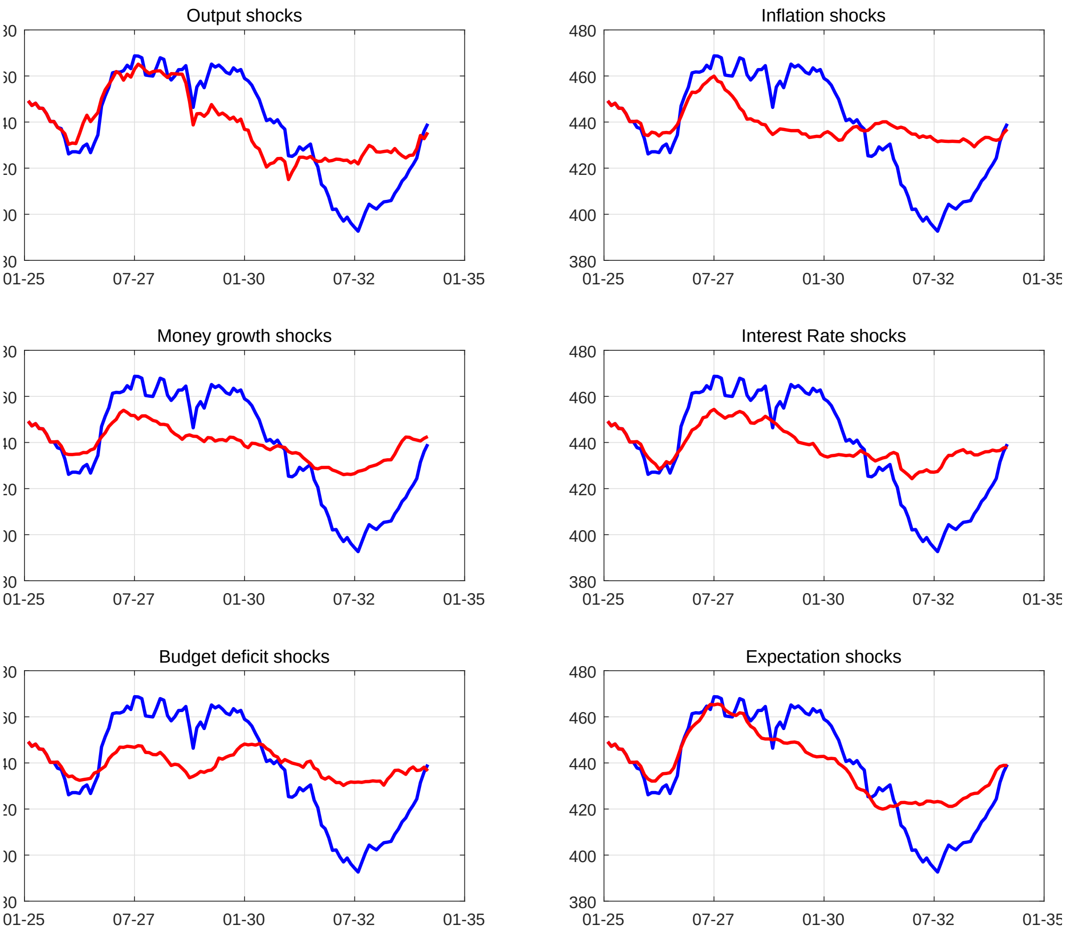
Notes: Implied inflation expectations from futures prices on wheat in percent per year. Inflation computed as the logged differences between futures prices at different time periods.

The results for this final exercise are shown in Figure 1.B.3 for maturities of two, four, and five months. We observe that during the summer of 1932, future prices of longer maturities were generally greater than prices of contracts that were soon to be settled, which would - at first glance - indicate that traders expected prices to increase during the following months. Notably, this is very likely the result of factors unrelated to inflationary policies. First, we observe fluctuations in future prices that indicate seasonal movements. During spring and summer, we would expect that traders forecast wheat prices to be higher than during the winter months. This indicates that traders expected higher prices before the harvest and lower prices after the fall, when wheat storages were filled. Therefore, the 1932 pattern in wheat futures is no different than it was before. Toward the end of the year, we observe declining expected wheat prices, inconsistent with upward shifts in expected inflation during the recovery. Second, we doubt whether futures prices can reveal informative information on inflation. Because futures prices of longer horizons already traded consistently above prices for shorter settlement dates in 1930 to 1931, we question if such prices actually revealed the expected deflation that should have been present in this period.

1.B.6 Robustness checks historical decomposition

Here we report additional results for the historical decomposition in the main text. We first checked whether we found different results when expected inflation is ordered last in the VAR, thus assuming that expected inflation reacts contemporaneously to all structural shocks. Results are presented in Figure 1.B.4, which confirms that the results presented in the main text are robust to this reordering.

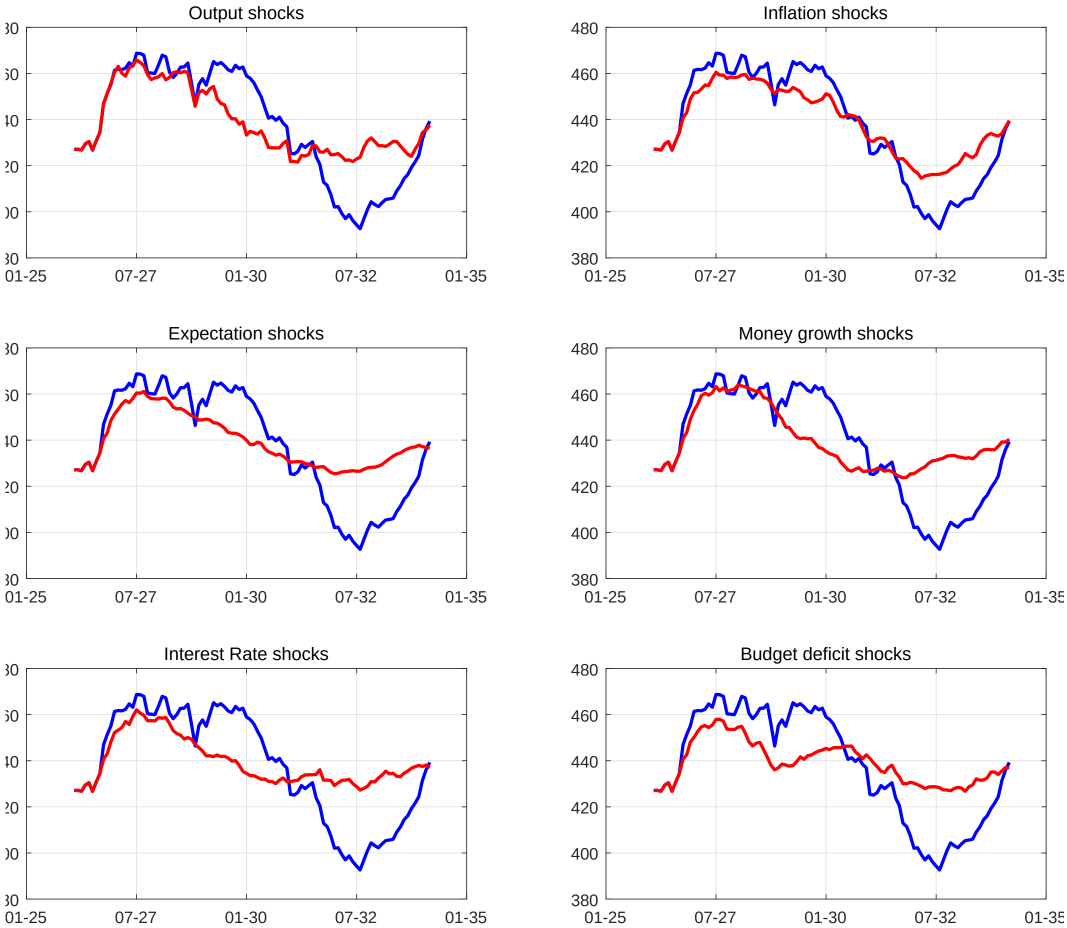
Figure 1.B.4: Historical decomposition of industrial production



Notes: Blue lines denote 100 times the log of the industrial production index. Red lines depict the series for industrial production that would have been obtained if only shocks to the respective variable occurred. In contrast to the main text, expected inflation is ordered last.

Secondly, we also tested whether using expected inflation from the real interest rate regressions instead of the news series forecast led to different results. As Figure 1.B.5 shows, this is not the case, and we observe results very similar to those presented in the main text.

Figure 1.B.5: Historical decomposition of industrial production



Notes: Blue lines denote 100 times the log of the industrial production index. Red lines depict the series for industrial production that would have been obtained if only shocks to the respective variable occurred. Expected inflation is based on the real interest rate regressions from the main text instead of the inflation news series forecasts. As in the main text, expected inflation is ordered third.

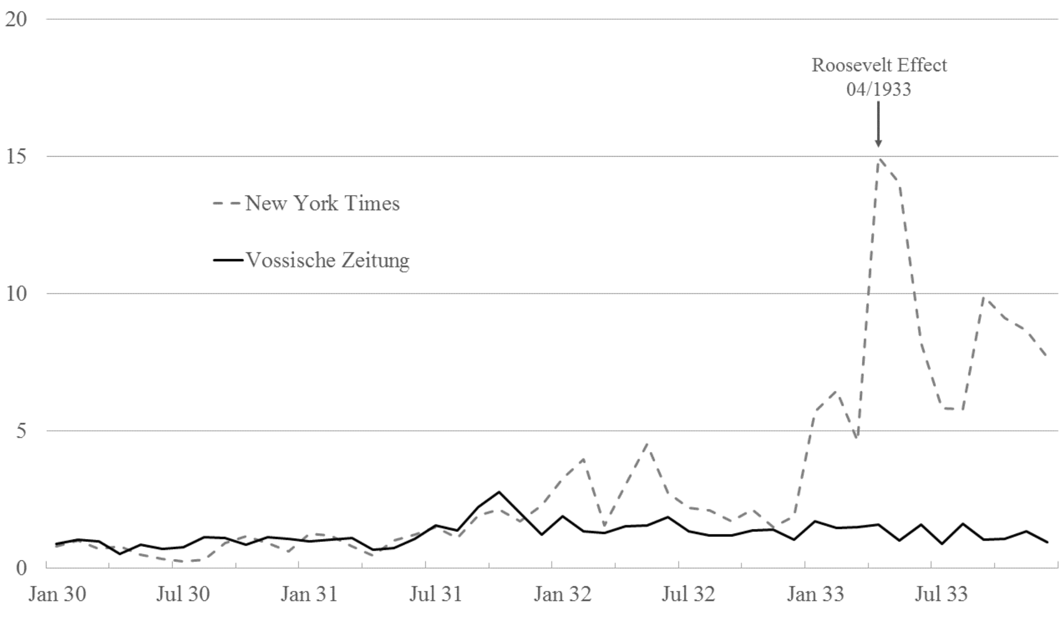
1.C News counts

1.C.1 Comparison of news counts in Germany and the United States

Figure 1.C.1 plots the daily average number of articles that contain the terms “inflation” or “inflationary” in *The New York Times* and comparable German expressions in *Vossische Zeitung* in each month over the period 1931 to 1934. In *The New York Times*, before 1933, approximately one to two articles per day contained one of the terms. In April 1933, this number spikes tenfold to a total of fifteen articles per day. Jalil and Rua (2016) surveyed the media account in the U.S. and found that many articles included in the figure are related to inflation. These articles demonstrate that the spike in news coverage in April 1933 indeed reflects a dramatic shift in inflation expectations in the U.S.

For Germany, the black line indicates that between 1930 and 1933, there was no increase of comparable magnitude. The series depicts the same data as in Figure 1.3 divided by the number of days in each month. Although both series have a similar magnitude before 1933, the numbers for the German newspaper vanish afterward when compared with the data of *The New York Times*.

Figure 1.C.1: Inflation mentions in *The New York Times* and *Vossische Zeitung* 1930 to 1933

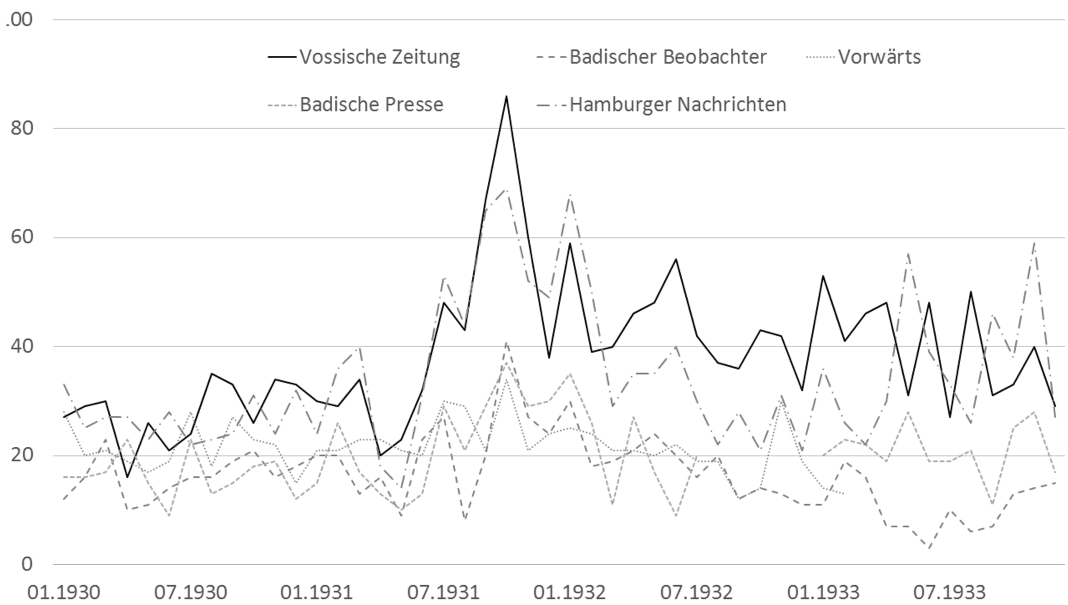


Notes: The daily average frequency in each month of articles that contain terms related to inflation in the *Vossische Zeitung* (black) and the *The New York Times* (grey-dashed) 1930 to 1934. We indicate April 1933 as the month in which the regime shift in inflation expectations occurred in the U.S. according to Jalil and Rua (2016). *Sources:* De Gruyter (2010), ProQuest (2004).

1.C.2 Additional news sources

As a robustness check for Figure 1.3, Figure 1.C.2 considers article counts from four additional newspapers plus *Vossische Zeitung*. We collected the total number of articles in each month from 1930 to 1933 that contain terms related to inflation (terms: "inflation", "teuerung", "preisheraufsetzung" or "reflation") in *Badischer Beobachter (1925/1934)* (a centrist newspaper from Karlsruhe in Southwest Germany), *Badische Presse (1925/1934)* (a more conservative newspaper also from Karlsruhe), *Hamburger Nachrichten (1925/1934)* (a Hamburg based national-socialist newspaper) and *Vorwärts* (socialist). Notably, the databases allow for searches on only total pages, whereas no classification at the article level is available. Therefore, we cannot exactly compare the series with the word counts illustrated in the main part of the paper.

Figure 1.C.2: Inflation series of additional newspapers 1925 to 1934



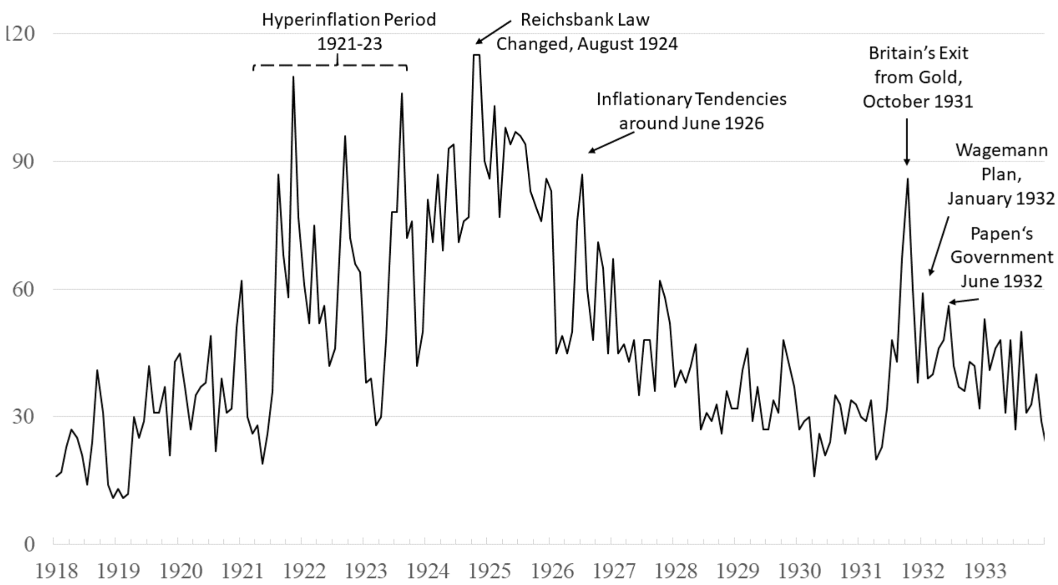
Notes: Inflation series of five newspapers. Each series denotes the number of mentions of inflationary terms in *Badische Presse*, *Badischer Beobachter*, *Hamburger Nachrichten*, *Vorwärts* plus *Vossische Zeitung* in each month 1930 to 1933. Note that the additional newspapers have gaps in the periods of observation due to missing values. The terms we searched for in the additional newspapers related to inflation were "inflation", "teuerung" "preisheraufsetzung" or "reflation".

We further recognized that the data quality of the text documents is lower than that in the case of *Vossische Zeitung*. Nonetheless, the contexts and situations with inflationary mentions were closely related to the news coverage in *Vossische Zeitung*. Although some spikes occur in different months, all five newspaper series clearly point to an increase in inflation mentions in October 1931 after Britain abandoned the gold standard and mention the other inflationary events we discuss in the narrative section of the paper such as the Wagemann plan in January 1932 and the change in government in June 1932.

1.C.3 The Origins of inflationary fears and longer-term inflationary news counts

Referring to the potential origins of German's fears of inflation, Figure 1.C.3 plots the series of inflationary terms in *Vossische Zeitung* over the prolonged period 1918 to 1934. The series spikes during the hyperinflation before 1923 and returns to high levels during the deflation after 1930.

Figure 1.C.3: Inflation mentions in the *Vossische Zeitung* 1918 to 1934



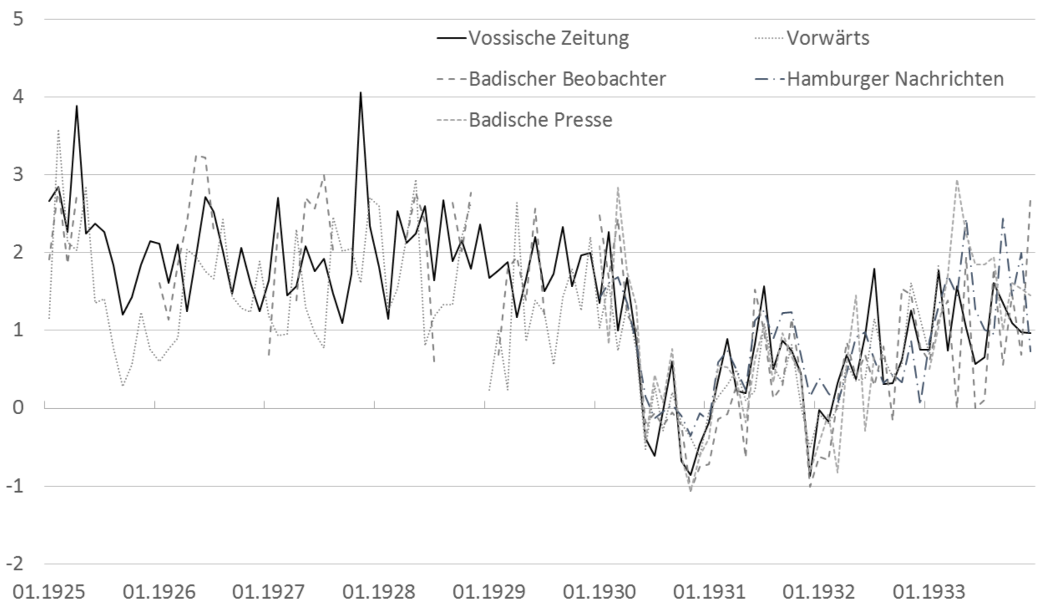
Notes: The monthly number of articles that contain terms related to inflation (black line: terms related to "inflation", "reflation" or "teuerung") in *Vossische Zeitung* 1918 to 1934. We indicate events with relation to inflation or fears of inflation as mentioned by Borchardt (1985).

In the figure, we indicate periods in which Borchardt (1985) noted debates with respect to inflation and fears of inflation in Germany in the 1920s and 1930s. Before the Depression, the article count spikes in the periods considered relevant by Borchardt: during Germany's hyperinflation 1921 to 1923, in response to a change in the Reichsbank law in August 1924 and in the economic boom period around June 1926. We further identified references to the indicated events in the news sources in the relevant months. Fears of another hyperinflation could have been one source of the ongoing debates on inflation in the 1930s.

1.C.4 Additional scaled inflation news series

Searching for inflationary terms raises the question of whether articles included in the list actually address deflation. The reverse is possible for articles including deflationary terms. A more convincing means to measure inflationary (positive or negative) news coverage is to scale the series as proposed by Binder (2016). For this purpose, and as shown in Subsection 1.4.5, we constructed a series of the logged difference between the number of articles containing inflation and the number of articles containing deflation per month. We illustrate this scaled measure together with the logged positive or negative inflation news counts in the first panel of Figure 1.8 of Subsection 1.4.5. The finding is that the scaled measure of media coverage on inflation takes up concerted actions to reduce prices in July 1930 and November 1930 or relevant laws and presidential decrees that included deflation in December 1931. Hence, the number of articles including deflation assumes some real effects of the discussed measures to reduce prices.

Figure 1.C.4: Scaled inflation series of additional newspapers 1925 to 1934



Notes: Scaled inflation series of five newspapers. Each series denotes the difference in the logged number of monthly articles that contain terms related to inflation and deflation in *Badische Presse*, *Badischer Beobachter*, *Hamburger Nachrichten*, *Vorwärts* plus *Vossische Zeitung* in each month 1925 to 1934. Note that the additional newspapers have later starting dates or gaps in the periods of observation due to missing values.

Comparing the scaled measure for a large number of additional newspapers as shown in Figure 1.C.4 reveals that the measure is robust to news outlets from different regions and political positions. For Figure 1.C.4, we used the word searches from the newspapers we already presented in Figure 1.C.2. We computed the measure by collecting additional deflationary terms: "deflation" (deflation), "preissenkung" (price reduction) and "preisherabsetzung" (price cut). The overall trends in the inflation measure are remarkably similar across all five newspapers over the period 1925 to 1933. However, the data quality differed considerably across the newspaper databases. We therefore hesitated to include all stated newspapers in the narrative and regression analysis of the main part of the paper.

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Chapter 2

Leaning against the wind and crisis risk *

Joint with Moritz Schularick and Felix Ward

2.1 Introduction

How should a central bank react when it observes that a potentially dangerous credit and asset price boom is under way? Can policymakers defuse rising financial stability risks by leaning against the wind and increasing interest rates?

Two prominent historical episodes delineate the issue that our paper speaks to. Consider the U.S. economy in 1928. Concerned about booming stock prices, a frenzy in commercial real estate markets, and substantial lending against both, the Federal Reserve increased policy rates from 3.5% to 6% between January 1928 and August 1929, surprising market participants. Most economic historians today think that these policy decisions, instead of bringing financial markets and credit growth back to more sustainable levels, played an important role in triggering the Great Depression (Bernanke, 2002; Eichengreen, 1992). Could the economy have avoided the financial crash had policy makers not raised interest rates to discourage what they perceived as rampant speculation in the stock market? Fast-forward 75 years. In the 2000s, U.S. policymakers decided to not lean against booming credit and housing markets. Instead, they stuck to a policy that was, by and large, consistent with flexible inflation targeting without taking financial stability considerations on

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board.[1] Financial imbalances continued to grow and erupted in the 2008 global financial crisis. What would have happened had the Federal Reserve raised interest rates to lean against the credit boom? Could the crash and the Great Recession have been avoided?

The two biggest financial crises in the past 100 years come with conflicting messages regarding the effectiveness of leaning against the wind (LAW) policies in safeguarding financial stability. Notwithstanding, current debates on the financial stability mandate of central banks often invoke one or the other episode to argue for or against leaning against the wind. Yet, which historical lesson is actually representative? The issue looms large for current thinking about monetary policy (Adrian and Liang, 2018; Stein, 2013; Svensson, 2017). It has become exceedingly clear how large the economic costs of financial crises are (Cerra and Saxena, 2008; Jordà et al., 2013). Moreover, recent research suggests that such financial boom states are detectable in real-time using quantity and price indicators (Richter et al., forthcoming) so that policymakers have the chance to intervene. Yet, what do we know about the effects of monetary policy changes on financial stability during financial booms? The answer so far is not much, other than inconclusive anecdotal evidence.

This paper aims to close this gap. We systematically study the available evidence for the state-dependent effects of monetary policy on financial stability. The state we condition on is a financial boom, defined as a large and sustained deviation of credit growth and real asset prices from trend. Conditional on being in such an (observable) boom state, we estimate how a monetary policy shock affects financial crisis probability and severity. We do so based on the near-universe of advanced economy financial crises since the 19th century.

Note that the question we are interested in is not how systematic LAW policy rules affect financial stability and the macroeconomy. Our focus is squarely on exogenous and unanticipated monetary policy actions that take place in a financial boom state, and our identification strategy speaks to those cases only. Our empirical analysis is based on a local projection instrumental variable (LP-IV) strategy that has recently been introduced by Jordà et al. (2019). The IV exploits a type of monetary policy variation that is not itself influenced by local economic conditions, namely, policy rate changes in small open economies with fixed exchange rates that are induced by the base economy.

For instance, in the early 1990s, Sweden witnessed a credit and house price boom. When the German Bundesbank surprised markets in December 1991 and raised its Lombard rate to 9.75% in response to inflationary pressures following German reunification, under the prevailing fixed exchange rate regime, it forced the

[1] There is some disagreement whether policy was too loose relative to an estimated Taylor rule with realized or expected values for inflation. See Bernanke (2010).

hand of the Swedish central bank too. At the time, the *New York Times (1991)* quoted a market economist: "This is the Bundesbank's way of showing they will use their power and independence without regard to the economic conditions in the rest of Europe." The Riksbank had to defend the exchange rate of the Swedish Krona vis-à-vis the German Mark. Following the Bundesbank, the Riksbank also increased its policy rate at a time when credit and housing markets in Sweden were booming. This episode provides us with a quasi-experiment for an exogenous change in monetary conditions at a time when credit and housing markets in Sweden were booming.

We bring this identification strategy to bear on a long-run dataset that spans 150 years and covers most advanced economies (Jordà et al., 2017), including dates of systemic financial crises. The dataset contains 1,525 country-year observations of countries whose currency is pegged to a base country's currency. Among those, we observe more than 170 credit boom episodes, of which 98 coincide with exogenous increases in base country policy rates. This rich dataset and the IV identification strategy allow us, for the first time, to zoom in on the causal effects of LAW policy—*increases* in policy rates *during booms* in credit and asset prices.

Our results are unambiguous in the sense that the estimates suggest that the effect of LAW policy on crisis risk has the opposite sign from what is often assumed. We show that a 1 percentage point (ppt) policy rate hike during a financial boom *increases* the risk of a financial crisis by about 10 ppts over a one-year horizon. Crisis risk remains elevated for about two years after the monetary shock before subsiding to its long-run average level. However, at no point in the five years following the policy rate increase do we find evidence for a reduction in crisis risk. The empirical evidence thus lends support to some of the worst fears about LAW policy—that it is more likely to trigger crises than prevent them (Bernanke and Gertler, 2000; Bernanke, 2002).

Although it heightens crisis risk in the near term, LAW policy could still be beneficial if it limits the economic costs of the crisis. We compare real GDP losses across financial crises that were preceded by different degrees of LAW prior to the start of the crisis, instrumenting the central bank's pre-crisis monetary policy stance. Our findings suggest that LAW policy does not systematically reduce crisis severity. In the five years after a financial crisis, real GDP falls by around 8% below trend, regardless of whether pre-crisis monetary policy was taking a leaning stance.

We corroborate these findings through a series of robustness checks. In particular, we examine alternative financial boom definitions, threats to the exclusion restriction, alternative financial crisis definitions, and differences between LAW interventions that take place early on versus late during financial booms. Throughout, the crisis trigger effect of LAW policy emerges as a robust feature of the data, whereas evidence for the crisis severity reduction effect remains elusive.

The empirical evidence brought together in this paper substantiates concerns that have been voiced by the opponents of LAW (Bernanke and Gertler, 2001; Gilchrist and Leahy, 2002; Svensson, 2017): contractionary monetary policy at best appears ineffective at addressing financial instability risks and at worst appears outright harmful (Bernanke and Gertler, 2000; Bernanke, 2002). Most existing studies of LAW policy focus on how monetary policy affects financial crisis risk and severity through its effect on credit growth (Bauer and Granziera, 2017; Svensson, 2017). The "credit-only" approach suggests that LAW policy decreases crisis risk and ameliorates crisis severity to the extent that it reins in pre-crisis credit growth. This approach underlies assessments of LAW policy (Ajello et al., 2016; Alpanda and Ueberfeldt, 2016; Gourio et al., 2018; Svensson, 2017). However, it is plausible that monetary policy affects financial stability also through other channels (e.g., through its effect on debt servicing costs, asset prices, and income). Our paper provides a direct causal estimate of the effects of monetary policy on financial stability that is agnostic with respect to the channels at work.

Theoretical studies have focused on monetary policy rules that incorporate LAW elements (Filardo and Rungcharoenkitkul, 2016; Juselius et al., 2017; Woodford, 2012). Such rules require the central bank to react to financial booms in a rule-based way. Currently, most central banks do not follow an explicit LAW policy rule. Any policy change in that direction would thus initially resemble a discretionary policy change until the commitment to the new policy regime has been credibly established (Svensson, 2016).[2] So while our paper speaks to the effects of state-dependent discretionary changes in monetary policy and not to the effects of systematic LAW, it can also inform the debates about the design and transition to systematic LAW policies.

The remainder of this paper is structured as follows. Section 2.2 introduces the data, and section 2.3 describes our empirical strategy. The results are presented in section 2.4. Section 2.5 concludes.

2.2 Data

Our main data source is the JST Macrohistory Database (Jordà et al., 2017, `http://www.macrohistory.net/data/`). It provides annual data on the real economy and the financial sector for 17 developed countries since 1870. The countries included in the sample are Australia, Belgium, Canada, Denmark, Finland, France, Germany,

[2] The relative infrequency of financial booms raises further questions about the extent to which central banks are able to credibly commit to a LAW policy rule, as well as the private sector's ability to systematically incorporate such a rule in its decision making.

Italy, Japan, the Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, the United Kingdom, and the United States.

To analyze how monetary policy affects financial crisis risk, we use the systemic financial crisis dummy defined by Schularick and Taylor (2012). This binary indicator is a narrative crisis measure that takes the value 1 in years in which a country experienced bank runs, bank defaults, forced mergers, or major public interventions in the financial sector. As a robustness check, we also consider the banking crisis dummies defined by Reinhart and Rogoff (2011) and Baron et al. (2018) (see Appendix 2.B). Regardless of the crisis indicator chosen, we obtain very similar results.

Our main explanatory variable of interest is the stance of monetary policy, which we measure as the change in nominal short-term interest rates. Other variables that enter our analysis also come from the JST Macrohistory Database.

2.2.1 Exchange rate regime, base countries, and capital controls

To construct the Trilemma IV for nominal short-term interest rate changes, we combine data on a country's exchange rate regime and capital account openness with data on interest rate changes in important base countries (see section 2.3). Our long-run exchange rate regime indicator comes from Jordà et al. (2019), relying on the work of Ilzetzki et al. (2019). We use a binary variable that classifies country-year observations as a peg (=1) if the exchange rate is fixed, or a float (=0) otherwise.

Besides knowing whether a country entertains a fixed or floating exchange rate regime, our empirical strategy requires that we define a base country with respect to which the exchange rate is fixed. This base country's interest rate—the base rate—is the primary source of variation in the Trilemma IV. In the definition of base countries, we follow Jordà et al. (2019). The U.K. is the base country prior to 1914. After 1945, the base is generally the U.S., with the exception of the ERM/EMS/Eurozone countries, for which Germany is treated as the base country. In the interwar years, we define a "gold rate," which is an average of U.K., U.S., and French short-term rates. Of the three countries, only those on gold are included in the average in any given year (see Obstfeld and Taylor, 2004). To capture the degree to which local interest rates are insulated from base country rates through capital controls, we make use of the capital mobility indicator by Quinn et al. (2011). Their index ranges from 0 to 100, with 0 indicating a low degree of capital mobility and 100 a high degree. We rescale this indicator to the 0-1 interval.

2.2.2 Financial boom indicators

To analyze the effect of LAW policy in the context in which it is usually considered (i.e. periods of rapidly expanding credit), we construct a binary indicator for credit booms. This boom indicator, $B_{i,t}$, takes a value of 1 when log real credit, $y_{i,t}$, is above its trend level, $\bar{y}_i$, and growing:

$$\mathrm{B}_{i,t} = I(y_{i,t} > \bar{y}_i \wedge \Delta y_{i,t} > 0). \tag{2.1}$$

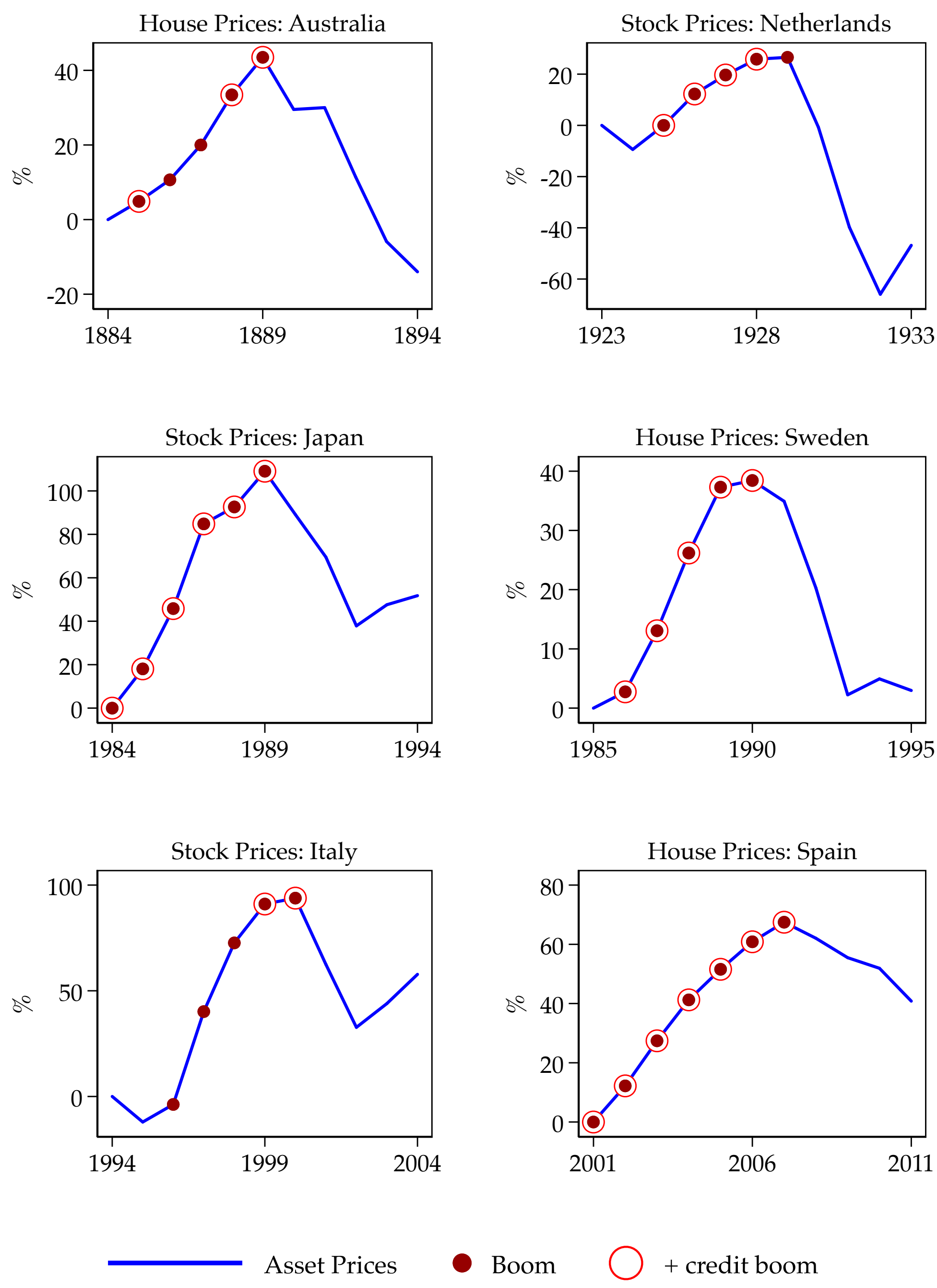
To obtain the cyclical component, we use the HP-filter with a smoothing parameter, λ, equal to 100 (Hodrick and Prescott, 1997). As a robustness check, we also consider the Christiano-Fitzgerald bandpass filter (Christiano and Fitzgerald, 2003), isolating fluctuations in the 2- to 16-year period range, as well as the novel non-parametric filtering method that has recently been proposed by Hamilton (2018). Results based on these alternative filtering methods are very similar to the baseline results we report in the main text (see Appendix 2.A). In all cases, the detrending is conducted in a one-sided fashion, so the results are relevant to policymakers who have to evaluate whether the economy is in a boom state or not in real time.

We also consider combined booms in credit and house prices, as well as in credit and stock prices. These two combined boom indicators take a value of 1 when credit and the respective asset price both fulfill condition 2.1. Finally, as a robustness check, we also partition boom episodes into early and late boom stages, in order to evaluate the claim that early LAW interventions are more effective at diffusing crisis risk. For this, the early boom stage is defined variably as either the first half of a boom episode, the first two years of a boom episode, or only the very first year of a boom episode.

Figure 2.1 looks at six historical time frames that are commonly associated with financial market booms. The solid blue lines depict the log of the financial variable of interest. The solid and hollow circles highlight years that the above-described procedure isolates as asset price booms and credit booms, respectively. Consistent with more general appraisals of financial market conditions, the 1880s in Australia, the 1980s in Sweden, and the 2000s in Spain are all identified as house price booms. The 1920s stock price boom in the Netherlands, the 1980s boom in Japan, as well as the 1990s dot-com boom in Italy are similarly well captured. In most years, these asset price booms were also underpinned by booms in credit.

Our sample contains a total of 255 credit booms. Of these, 142 coincided with booms in house prices and 168 with booms in stock prices. Importantly for our identification strategy, our sample contains 171 credit boom episodes in countries with a fixed exchange rate. The respective numbers for the joint credit+house price, and credit+stock price booms are 100 and 113. During 98 of the pegs' credit boom episodes, the pegged country was exposed to a policy rate hike in the base country and thus participated in the quasi-experiment. The respective numbers for the joint credit+house price, and credit+stock price booms are 67 and 75.

Figure 2.1: Asset prices and boom periods



Notes: Blue lines – log of real asset prices in % deviation from asset price level 5 years prior to turning point. Solid circles – asset price boom years. Hollow circles – credit boom years.

2.3 Empirical strategy

To identify the effects of monetary policy on crisis risk, we apply the Trilemma instrumental variable (IV) strategy pioneered by Jordà et al. (2019). The reasoning behind the Trilemma IV strategy is as follows. When a country pegs its exchange rate to a base country's currency, the local interest rate from then on has to match that of the base country. To see why, consider the peg country setting its interest rate below that of the base country. This will lead to unsustainable capital outflows, as capital seeks to obtain the highest return. Vice versa, if the peg sets its interest rate too high, this results in unsustainable capital outflows. Hence, under perfect capital mobility, peg country interest rates have to move in sync with base country interest rates. Furthermore, because base country interest rate changes are determined only by base country economic conditions, their variation is exogenous to the economic conditions in the peg countries.

We also subtract predicted base country interest rate changes, $\Delta\hat{r}_b$, from actual rate changes, Δr_b, to isolate unpredictable movements in base country interest rates. The Trilemma IV is constructed using this unpredictable component to prevent the IV from conveying information about base rate changes that could have been anticipated by the peg's households and firms. As in Jordà et al. (2019), we use the first lags of the following base country variables to predict base country interest rate changes: the growth rates of GDP, consumption, investment, stock prices, and credit (all CPI deflated), as well as changes in nominal long-term interest rates, nominal short-term interest rates, the CPI inflation rate, and the current account-to-GDP ratio.

The Trilemma IV, z, for local policy rate changes, Δr, is thus defined as

$$z_{i,t} \equiv \left(\Delta r_{b(i,t),t} - \Delta\hat{r}_{b(i,t),t}\right) \times PEG_{i,t} \times PEG_{i,t-1} \times KOPEN_{i,t}. \tag{2.2}$$

where i and t are the country and year indices, $b(i,t)$ denotes country i's base country in year t, and $PEG_{i,t}$ is the exchange rate regime dummy that indicates whether a country's exchange rate is fixed or floating with respect to the base b. The lagged dummy ensures that the instrument includes only well-established pegs that have lasted for at least two years, excluding the most fleeting and possibly incidental single-year pegs. $KOPEN_{i,t}$ is the rescaled financial openness indicator. Jordà et al. (2019) show that the Trilemma IV, $z_{i,t}$, is closely aligned with changes in pegs' domestic short-term rates and is thus clearly relevant. In our sample, the instrument exhibits a highly significant slope coefficient of 0.6 over the full sample and 0.75 for the post-World War II sample.

To trace the effect of a +1 ppt increase in policy rates on crisis risk, we estimate impulse response functions (IRFs) through local projections. More particularly, the sequence of fixed effects models we estimate represent a sequence of linear crisis probability models through which we can assess how monetary policy affects crisis risk over a five-year horizon, $h = 0, ..., 5$:

$$C_{i,t+h} = \alpha_{i,h} + \beta_h^{IV} \Delta r_{i,t} + \sum_{l=0}^{L} \mathbf{\Gamma}_{h,l} \boldsymbol{X}_{i,t-l} + \epsilon_{i,t+h}, \tag{2.3}$$

where $\alpha_{i,h}$ denote country fixed effects, $\Delta r_{i,t}$ is instrumented by $z_{i,t}$, and $\boldsymbol{X}_{i,t}$ contains additional control variables. The dependent variable is a dummy, $C_{i,t}$, that takes the value 1 if a financial crisis occurs in country i in year t or in any of the following two years, $t+1$, $t+2$. This definition reflects that while it is notoriously hard to predict the exact crisis year, it is possible to predict whether an economy enters a *danger zone* in which financial crises are more likely to occur (e.g., Kaminsky and Reinhart, 1999; Ward, 2017). The coeffcients $\{\beta_h^{IV}\}_{h=0}^{H}$ trace out the response of crisis risk to a +1 ppt increase in monetary policy rates. We translate the three-year crisis probability IRFs into annual crisis probability IRFs using $\widehat{P}^{\text{annual}} = 1 - (1 - \widehat{P}^{\text{three years}})^3$.

We include a rich set of control variables, $\boldsymbol{X}_{i,t}$. In particular, we include four lags of the following variables: per capita GDP growth, consumption growth, investment growth (all in real terms), CPI inflation, a measure of world GDP growth as in Jordà et al. (2019), changes in short-term and long-term interest rates, growth rates of real stock prices, real house prices, real bank loans, the current account-to-GDP ratio, and the binary crisis dummy on which $C_{i,t}$ is based. Note that, except for the crisis dummy, we include the time t realizations of all control variables. Thus, we take a conservative stance with respect to the contemporaneous response of the dependent variable to monetary policy, effectively attributing as much as possible of that response to contemporaneous variation in the control variables and not the policy rate change. As a robustness check, we also apply the spillover correction proposed by Jordà et al. (2019), which immunizes our results against potential violations of the exclusion restriction brought about by international goods and financial market spillovers (see Appendix 2.C).

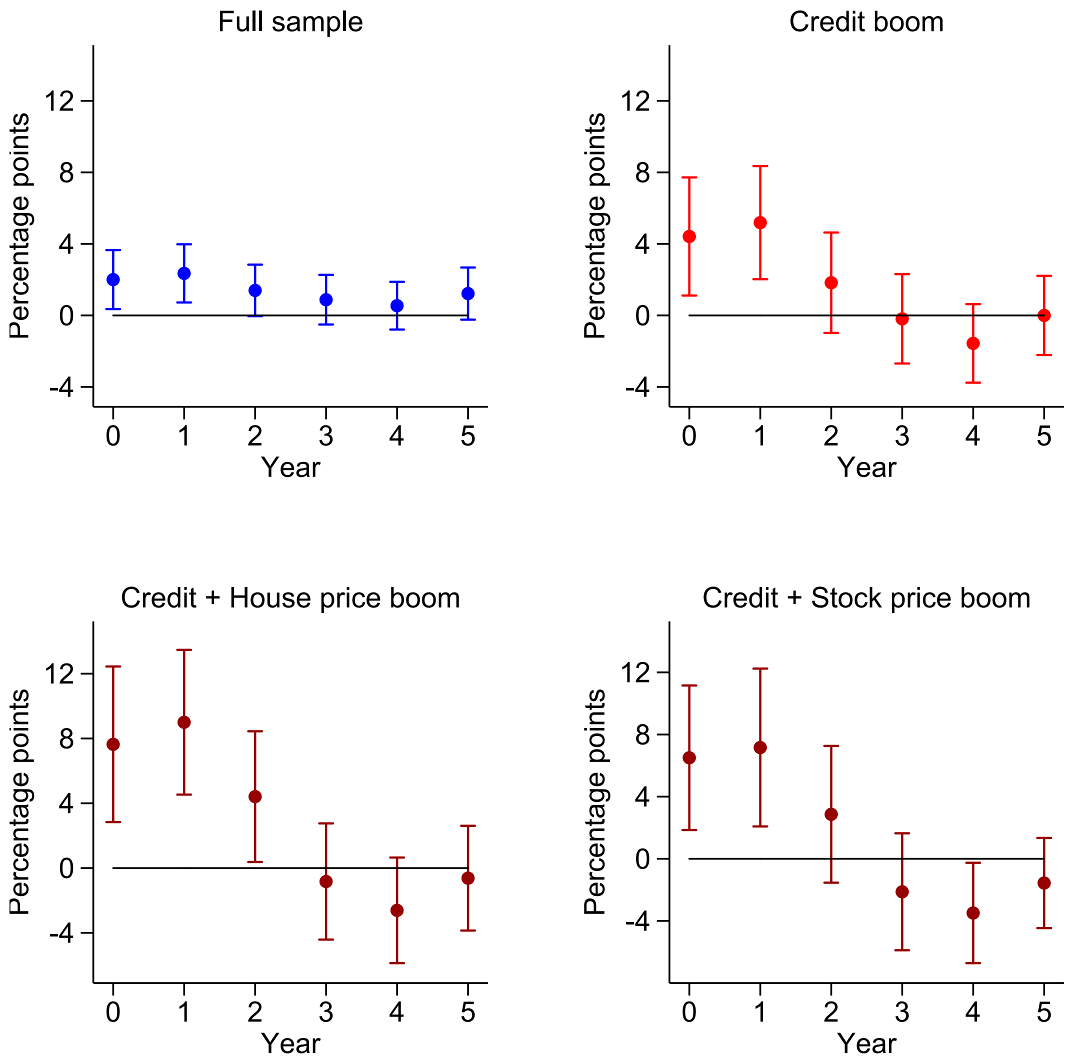
2.4 Results

Can contractionary monetary policy diffuse crisis risk? This section presents the answer provided by our empirical results. We begin by reporting the full sample results and then narrow down to LAW policy as conventionally defined—as policy rate hikes against the backdrop of financial booms. In a second step, we investigate the effects of LAW on crisis severity.

2.4.1 The effect of LAW on crisis probability

The full sample results in the top left panel of Figure 2.2 suggest that interest rate hikes increase crisis risk in the near term. More precisely, a +1 ppt policy rate hike increases crisis risk by 2 ppts on impact, as well as in the following year. The size of this effect is substantial, given that average annual crisis risk in the full sample is 3.4%.

Figure 2.2: Financial crisis risk responses



Notes: Change in the annual crisis probability following a monetary policy shock. 95% confidence bands.

Next, we consider the effect of contractionary monetary policy for subsamples of financial booms. Can LAW policy rein in crisis risk against the backdrop of soaring credit aggregates and asset prices? The top right panel in Figure 2.2 shows our credit boom subsample results, which suggest that an interest rate hike during credit booms has a particularly adverse effect on crisis risk. Crisis risk increases by 4 ppts on impact, as well as in the year following the LAW policy. Taking into account that crisis risk is already elevated during credit booms, a +1 ppt interest rate increase raises annual crisis risk from 4.8% to around 10%.

The subsample results for combined booms in credit and asset prices point in the same direction. A discretionary +1 ppt increase in interest rates, aimed at reining in equity or house price booms, increases crisis risk by 6 to 8 ppts for up to two years. Given that average crisis risk is already 5.2% in the credit + house price boom subsample and 4.7% in the credit + stock price boom subsample, the LAW policy raises crisis risk above 10% in the short term.

These findings lend empirical substance to the concern that LAW policies might provoke financial crises rather than prevent them. We find little evidence to support the notion that LAW policy may pay off in the form of lower crisis risk in the medium term. The only significantly negative effect of LAW policy on crisis risk that we can document occurs in year 4 after the interest rate hike in the combined credit + stock price boom subsample.

Policy rate hikes versus cuts

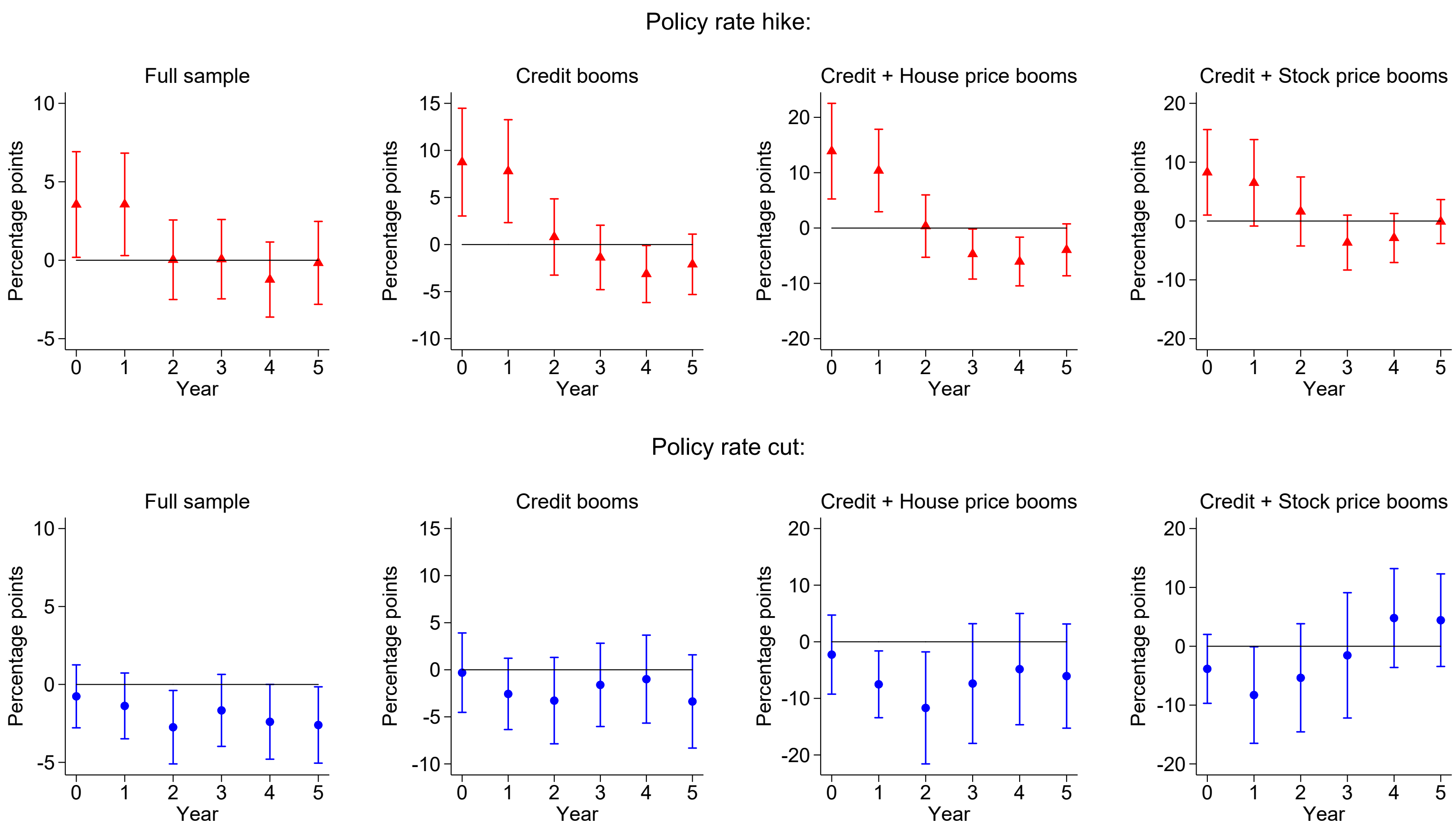
Resent research suggests that policy rate increases have stronger effects on the economy than policy rate decreases (Angrist et al., 2017; Tenreyro and Thwaites, 2016). This finding is relevant for LAW policy, which is commonly defined asymmetrically—as policy rate hikes during booms. Does crisis risk respond differently to policy rate hikes and cuts?

To answer this question, we augment our baseline specification (eq. 2.3) with an interaction term that separates positive changes in the instrumented policy rate from negative ones,

$$C_{i,t+h} = \alpha_{i,h} + \beta_h^{IV} \Delta r_{i,t} + \gamma_h^{IV} \Delta r_{i,t} \cdot hike_{i,t} + \sum_{l=0}^{L} \boldsymbol{\Gamma}_{h,l} \boldsymbol{X}_{i,t-l} + \epsilon_{i,t+h}, \tag{2.4}$$

where $hike_{i,t}$ is a dummy that takes the value 1 for policy rate hikes and 0 otherwise. This specification allows us to search for asymmetries in the response of crisis risk: $\{\beta_h^{IV} + \gamma_h^{IV}\}_{h=0}^{H}$ traces out the crisis risk response to policy rate increases, whereas $\{\beta_h^{IV}\}_{h=0}^{H}$ shows the same response for policy rate decreases.

Figure 2.3: Rate hikes versus rate cuts and crisis risk



Notes: Change in the annual crisis probability following a 1 ppt policy rate hike/cut. 95% confidence bands.

Figure 2.3 shows the asymmetry results for the full sample, as well as the three financial boom subsamples. The immediate increase in crisis risk after a policy rate hike stands out, regardless of subsample. For the full sample, a +1 ppt rate hike increases short-term crisis risk by 3.6 ppts—almost two times the effect size of the symmetric specification. For the financial boom subsamples, a full percentage point rate hike increases annual crisis risk by 8 to 14 ppts. Evidence for medium-term crisis risk reduction again is scant.[3]

Policy rate cuts tend to be followed by decreases in crisis risk. However, this crisis risk reduction effect tends to be less immediate than in the case of contractionary rate hikes. Only with a lag of one to two years does crisis risk decline significantly. For the full sample, the crisis risk reduction effect amounts to 2.7 ppts after two years. In the joint boom subsamples, crisis risk falls more substantially in a shorter amount of time.[4] A (pointwise) Wald test for equality of the rate hike and cut responses, however, indicates that the above-mentioned asymmetries in crisis risk responses are rarely statistically significant.[5]

The finding that policy rate hikes give rise to especially large increases in financial crisis risk strengthens the earlier contraindication result against LAW policy. This appears particularly pertinent against the backdrop of financial booms—precisely when LAW policy moves are usually considered.

Early versus late interventions

Maybe rate hikes trigger financial crises only when they are administered too late in the boom. By contrast, the same rate hike might diffuse crisis risk when administered early on in the boom.[6]

To empirically test this idea, we extend our baseline specification by an interaction term that allows the effects of LAW policy to differ for early and late interventions:

$$C_{i,t+h} = \alpha_{i,h} + \beta_h^{IV} \Delta r_{i,t} + \gamma_h^{IV} \Delta r_{i,t} \cdot early_{i,t} + \sum_{l=0}^{L} \mathbf{\Gamma}_{h,l} \boldsymbol{X}_{i,t-l} + \epsilon_{i,t+h}, \tag{2.5}$$

where all terms are defined as before, the policy rate changes $\Delta r_{i,t}$ are again instrumented by the Trilemma IV, and $early_{i,t}$ is a dummy variable that takes the value 1

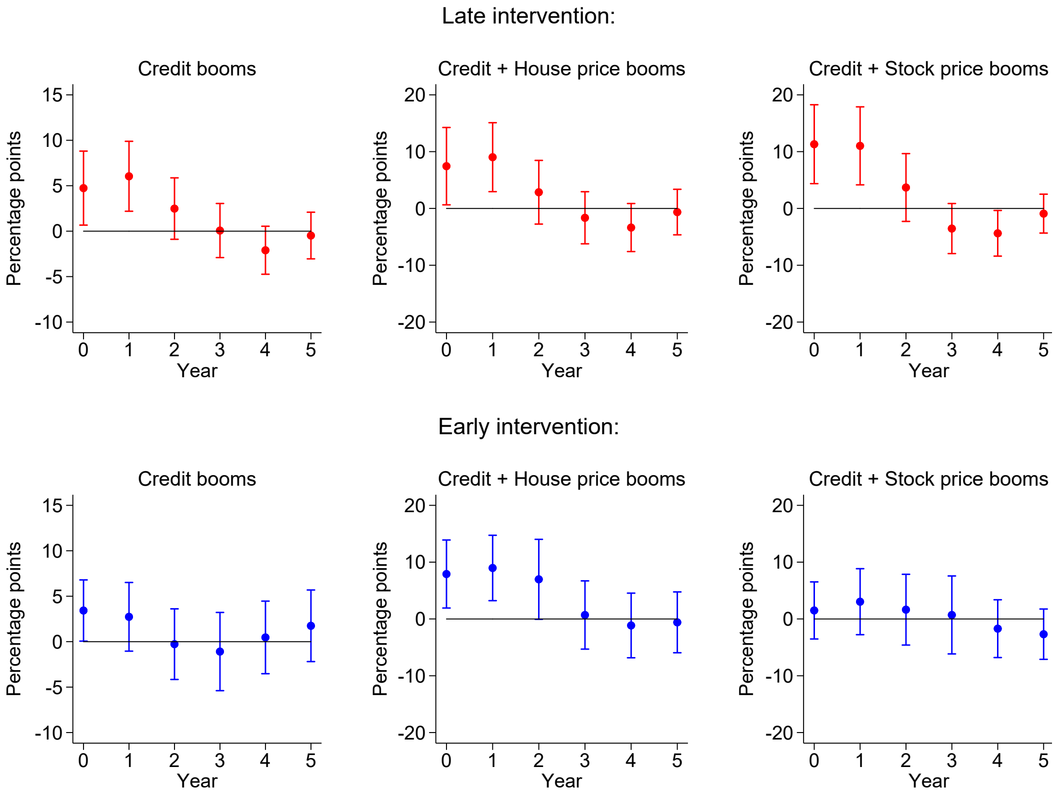
[3] Only for the credit and the credit + house price boom subsamples do we find isolated coefficient estimates that are in line with medium-term crisis reduction effects. The absolute size of these negative coefficients, however, is small compared to the initial crisis trigger effect.

[4] Additional results reported in Appendix 2.A confirm the robustness of these findings using the Christiano-Fitzgerald bandpass filter (Christiano and Fitzgerald, 2003), and the Hamilton filter (Hamilton, 2018) to define financial boom episodes.

[5] Only sporadically, in the credit boom and credit + stock price boom subsamples, does the Wald test reject equality of the rate hike and rate cut responses in the short run (90% confidence level).

[6] The limit cycle framework by Beaudry et al. (2015) allows for a formalization of this notion.

Figure 2.4: Early versus late interventions and crisis risk



Notes: Change in the annual crisis probability following a 1 ppt policy rate hike. 95% confidence bands.

in the first year of a boom episode. We also considered other definitions of "early," such as the first two years of a boom or the first half of a boom. The results for these alternative partitions between early and late boom interventions are very similar to the baseline results reported here (see Appendix 2.C).

Figure 2.4 shows how crisis risk responds to a 1 ppt increase in policy rates, early and late during a financial boom. In no case do we find evidence for the notion that early interventions can lower financial crisis risk. For the credit boom subsample, the early and late intervention IRFs both exhibit a crisis trigger effect, though the mean estimate suggests that it is smaller for early interventions. A Wald test for equality of the late and early intervention IRFs, however, cannot reject the null hypothesis that both IRFs are equal.

For the credit + house price subsample early and late interventions have very similar effects throughout. Only for the credit + stock price subsample do we find evidence that early interventions are significantly less harmful than late interventions. Even in that case, however, rate hikes do not lower crisis risk—they just do not appear to trigger crises.

In sum, while early interventions may be somewhat less harmful than late interventions, they do not appear to systematically lower crisis risk. At best, early interventions leave crisis risk unaffected. At worst, early boom interventions appear to be just as potent in triggering financial crises as late boom interventions are.

2.4.2 The effect of LAW on crisis severity

In spite of the crisis trigger effect, LAW policy could still be beneficial if, by causing a small crisis now, it prevents a much bigger crisis later on. In other words, by hindering booms from proceeding unchecked, LAW policy might limit the fallout from the subsequent bust.

We investigate this hypothesis by looking at whether LAW reduces the real GDP loss associated with financial crises. To do this, we first characterize the degree of leaning over a one-year, three-year, and five-year horizon as the cumulative sum of policy rate changes over the same time period, $\Delta^K r_{i,t-1} \equiv \sum_{k=1}^{K} \Delta r_{i,t-k}, K = 1, 3, 5$. How leaning affects financial crisis severity is then estimated on the basis of the following local projections:

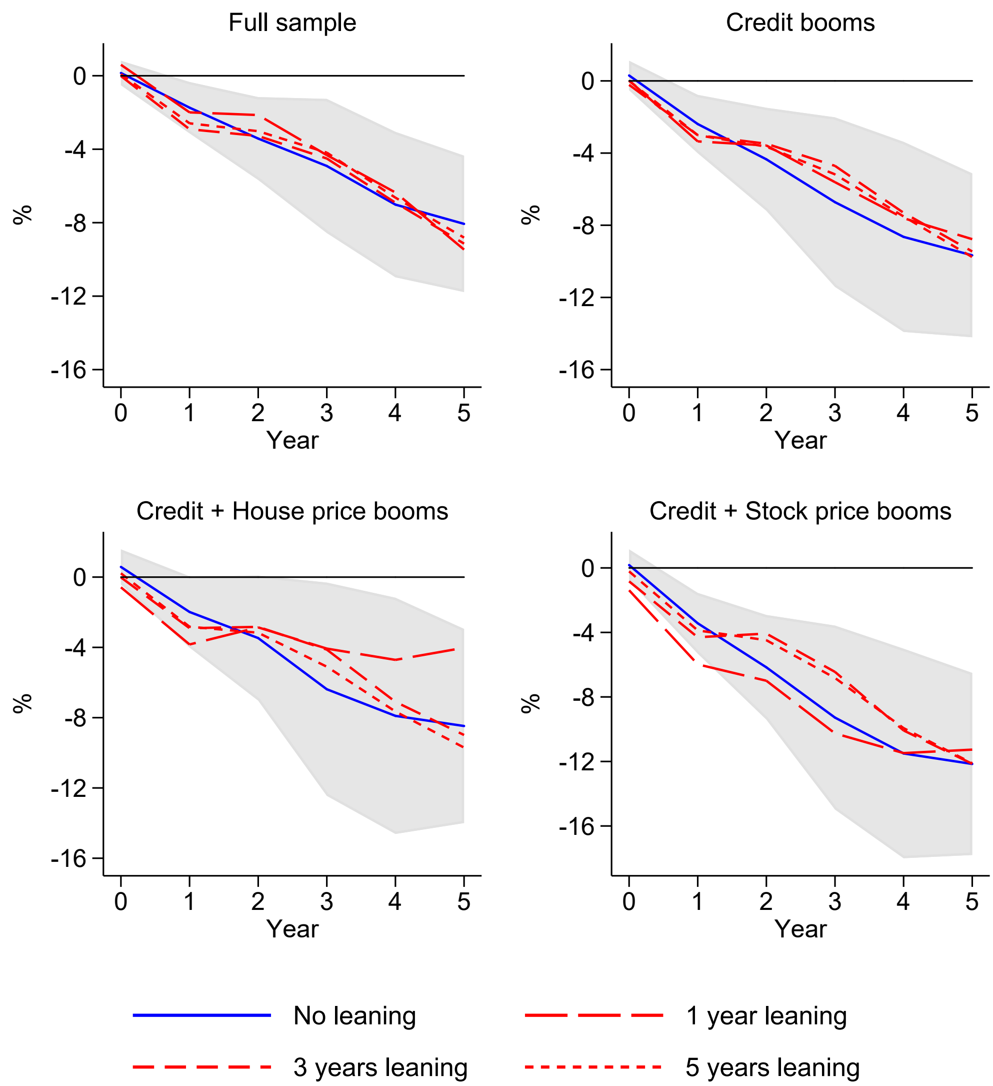
$$\Delta^h y_{i,t+h} = \alpha_{i,h} + \beta_h C_{i,t} + \gamma_h^{IV} C_{i,t} * \Delta^K r_{i,t-1} + \sum_{l=0}^{L} \mathbf{\Gamma}_{h,l} \boldsymbol{X}_{i,t-l} + \epsilon_{i,t+h}, \qquad (2.6)$$

where $\Delta^h y_{i,t+h}$ denotes the cumulative h-year change in real GDP, $\Delta^K r_{i,t-1}$ denotes the central bank's leaning stance in the years leading up to the crisis, and all other terms are defined as before.[7] The $\Delta^K r_{i,t-1}$ are again instrumented by the equivalent expression based on the Trilemma IV, and local projections are estimated separately for each of the three leaning periods, $K = 1, 3, 5$. $\{\beta_h\}_{h=0}^{5}$ describes how the real GDP path after a financial crisis deviates from its usual path after non-crisis years, and $\{\gamma_h^{IV}\}_{h=0}^{5}$ reveals how the real GDP path after a financial crisis is affected by leaning. If a leaning policy systematically lowers crisis severity, this should be indicated by estimates of γ_h^{IV} that are larger than zero. As a consequence, the path traced out by $\{\beta_h + \gamma_h^{IV}\}_{h=0}^{5}$ should lie above the path described by $\{\beta_h\}_{h=0}^{5}$.

Figure 2.5 shows the results for the full sample and the three boom subsamples. Our findings do not lend support to the idea that a leaning policy lowers crisis severity. Real GDP falls by around 8% regardless of whether monetary policy took a leaning stance or not. Only in the joint credit + house price boom subsample does the mean estimate indicate that crisis severity might be lower following a one-year leaning period. Large standard errors, however, render this difference statistically

[7] Note that the control vector $\boldsymbol{X}_{i,t}$ includes contemporaneous interest rate changes, which in contrast to the earlier specifications no longer receives separate mentioning here. $\boldsymbol{X}_{i,t}$ also contains the non-interacted leaning term $\Delta^K r_{i,t-1}$ to control for the effects of the central bank's policy stance in non-crisis years.

Figure 2.5: LAW and crisis severity



Notes: Real GDP loss after a crisis, depending on whether monetary policy was leaning against the wind or not. 95% confidence bands.

insignificant. We confirm the finding that leaning policy cannot be relied upon to lessen crisis severity on the basis of the banking crisis dummies defined by Reinhart and Rogoff (2011) and Baron et al. (2018) (see Appendix 2.B).

2.5 Conclusion

Whether conventional monetary policy should be applied to address financial stability risks is a long-standing question in macroeconomics. In this paper, we present the most comprehensive empirical analysis of LAW episodes in modern economic history. Our findings lend support to the concern that contractionary monetary policy increases financial crisis risk rather than reducing it. A policy rate hike increases crisis risk for up to two years, with little evidence that this short-term effect is compensated by either lower crisis risk in the medium-term or a reduction in crisis severity.

Our results add a new perspective to the current debate about whether macroprudential policy or monetary policy is better suited to address the buildup of financial fragilities. While monetary policy "gets into all the cracks" (Adrian and Liang, 2018; Stein, 2013), the empirical evidence points to severe side effects of discretionary LAW policies.

APPENDICES

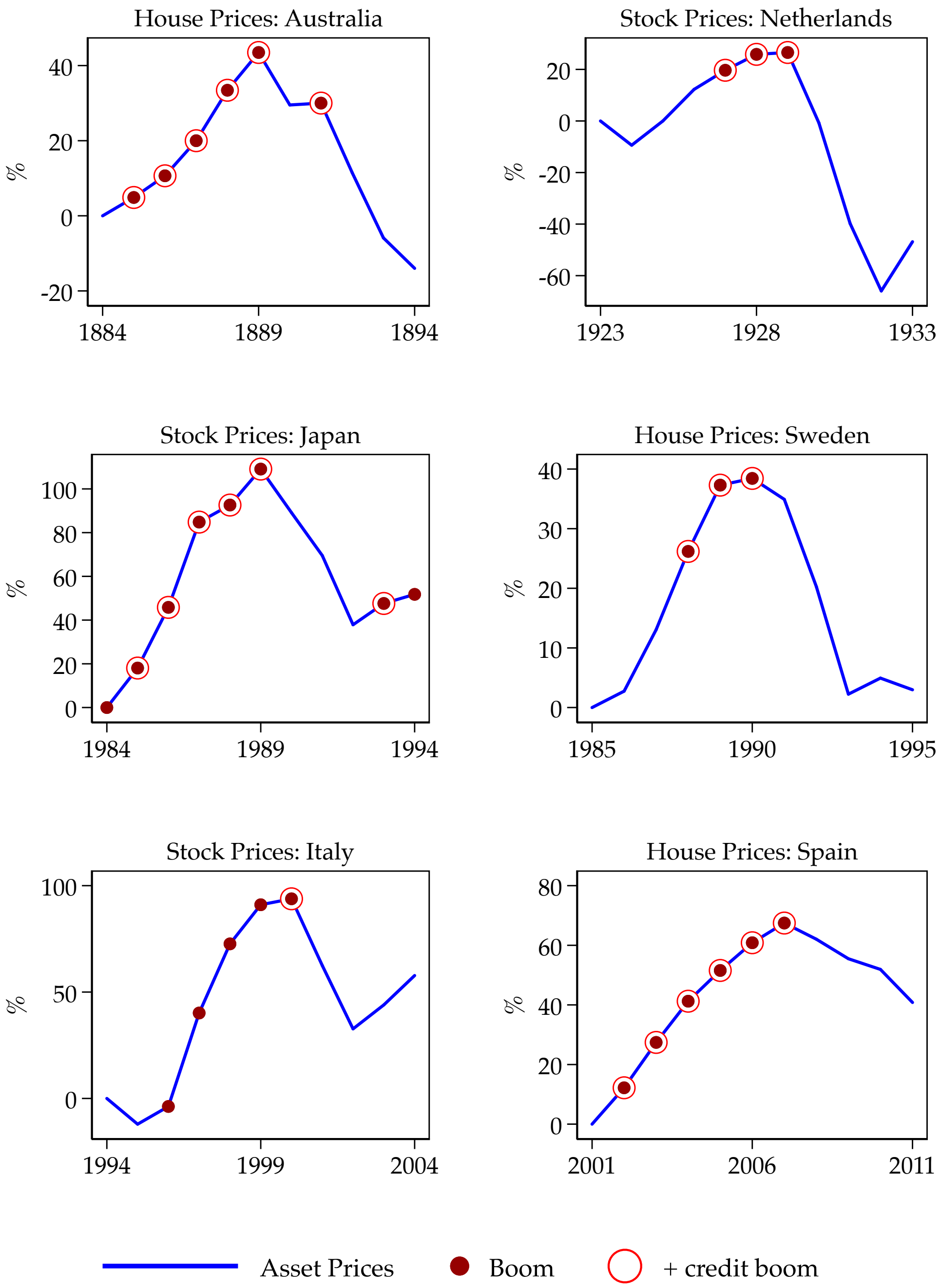
2.A Alternative time series filters

The main results in the paper use the one-sided HP filter to obtain cyclical components of credit and asset price time series to date boom periods. Here we investigate the robustness of our results with regard to that choice by using two alternative time series filters. Specifically, as the first alternative, we use the Christiano-Fitzgerald bandpass filter (Christiano and Fitzgerald, 2003), which we specify such that fluctuations in the 2- to 32-year period range are isolated. The results for the boom phases are shown in Figure 2.A.1, while the crisis risk responses and the asymmetric responses are shown in Figure 2.A.3, Figure 2.A.5, and Figure 2.A.7.

As the second alternative, we employ the novel non-parametric filtering method that has recently been proposed by Hamilton (2018), in which the cyclical component of a time series is defined as the residuals from an OLS regression of future values of the time series on a constant and its own lags. Results using this filter for boom phases are shown in Figure 2.A.2, and the results for the crisis risk responses and the asymmetric responses are shown in Figure 2.A.4, Figure 2.A.6, and Figure 2.A.8.

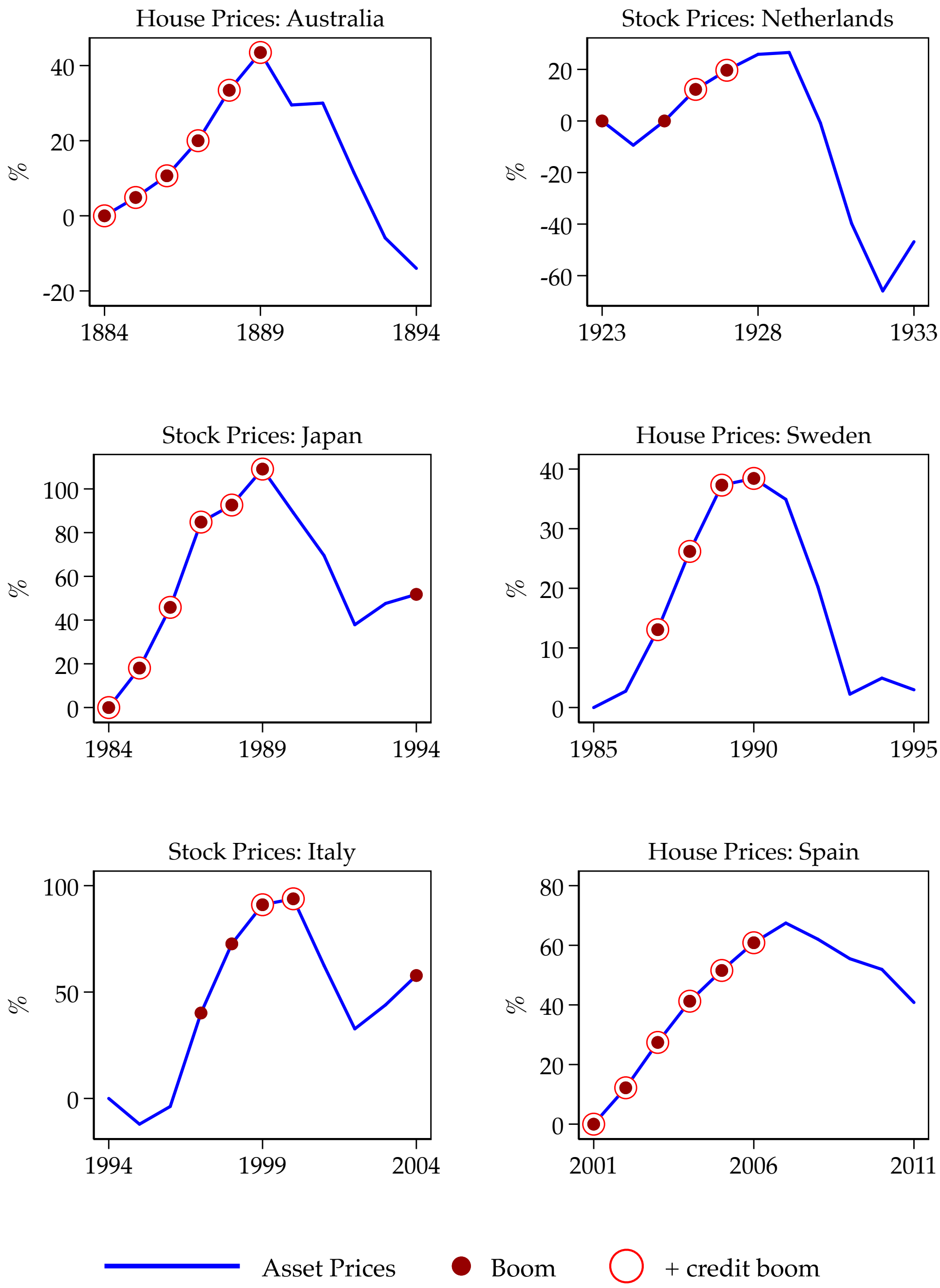
For both alternative filters, we find very similar classifications of boom phases and also similar results for the baseline empirical specification as well as for the asymmetric specifications.

Figure 2.A.1: Asset prices and boom periods—Christiano-Fitzgerald filter



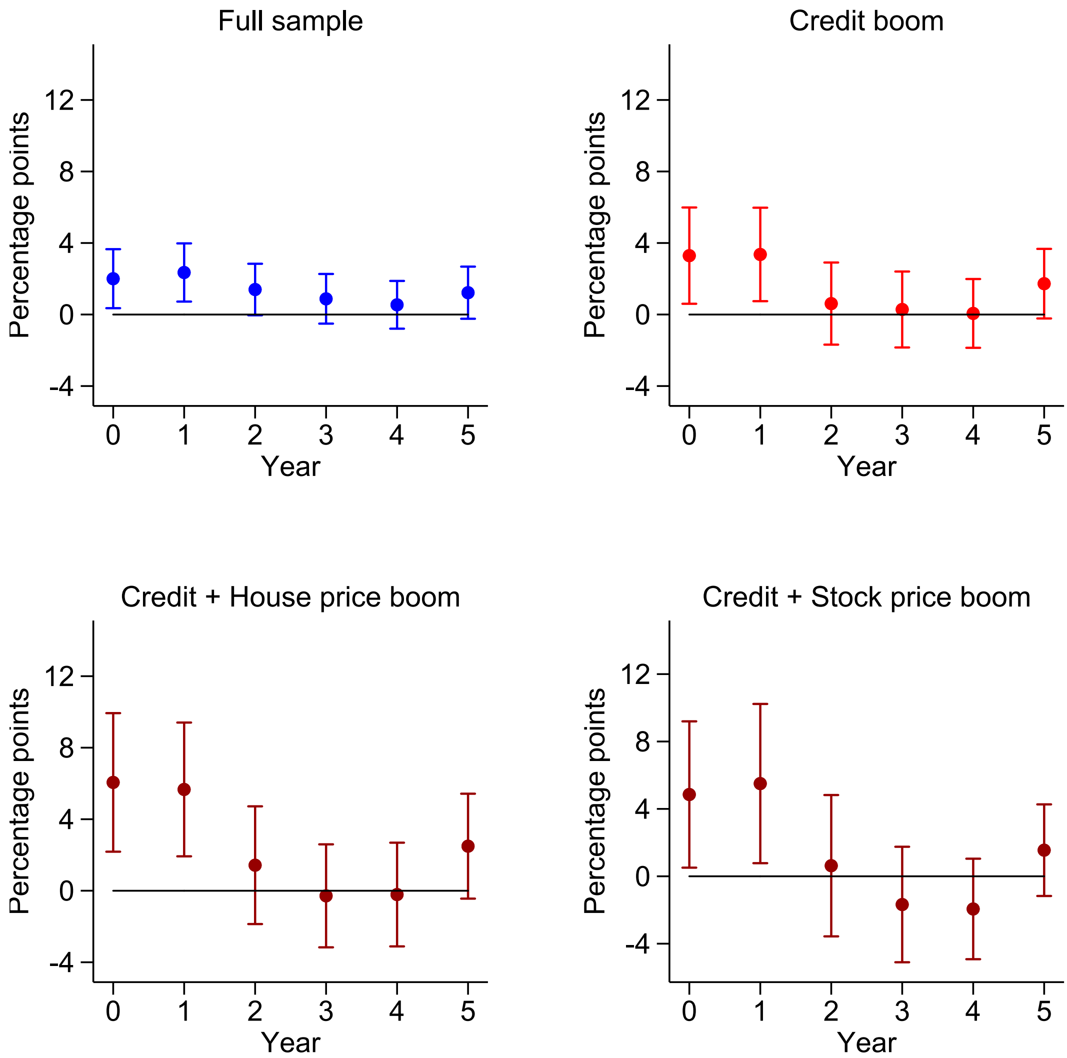
Notes: Blue lines denote the log of real asset prices, red circles denote the years identified as boom periods. The time series for asset prices are rescaled to start at 0 for each window. Boom episodes defined on the basis of CF-filtered series (2- to 32-period range).

Figure 2.A.2: Asset prices and boom periods—Hamilton filter



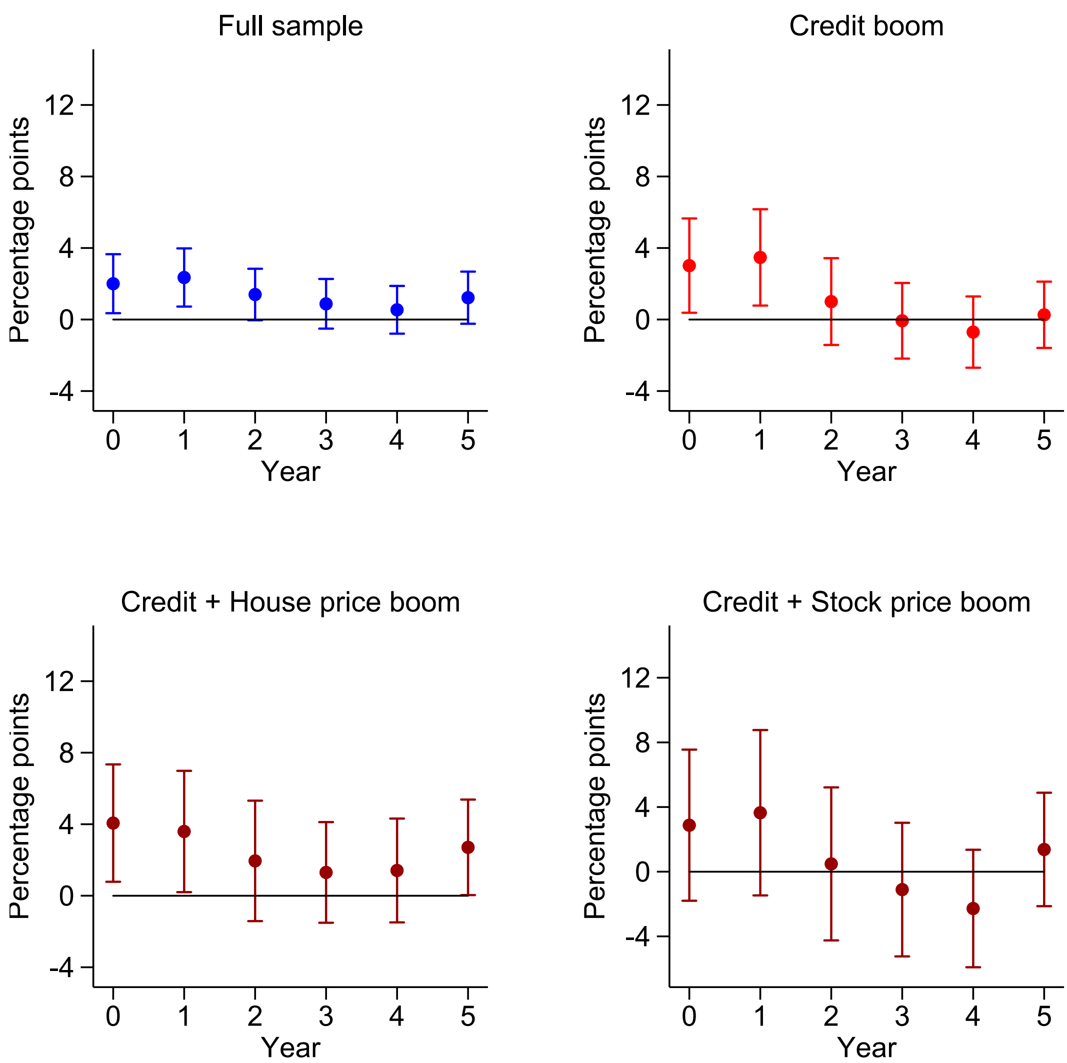
Notes: Blue lines denote the log of real asset prices, red circles denote the years identified as boom periods. The time series for asset prices are rescaled to start at 0 for each window. Boom episodes defined on the basis of Hamilton-filtered series (lags 3 to 6).

Figure 2.A.3: Crisis risk response—Christiano-Fitzgerald filter



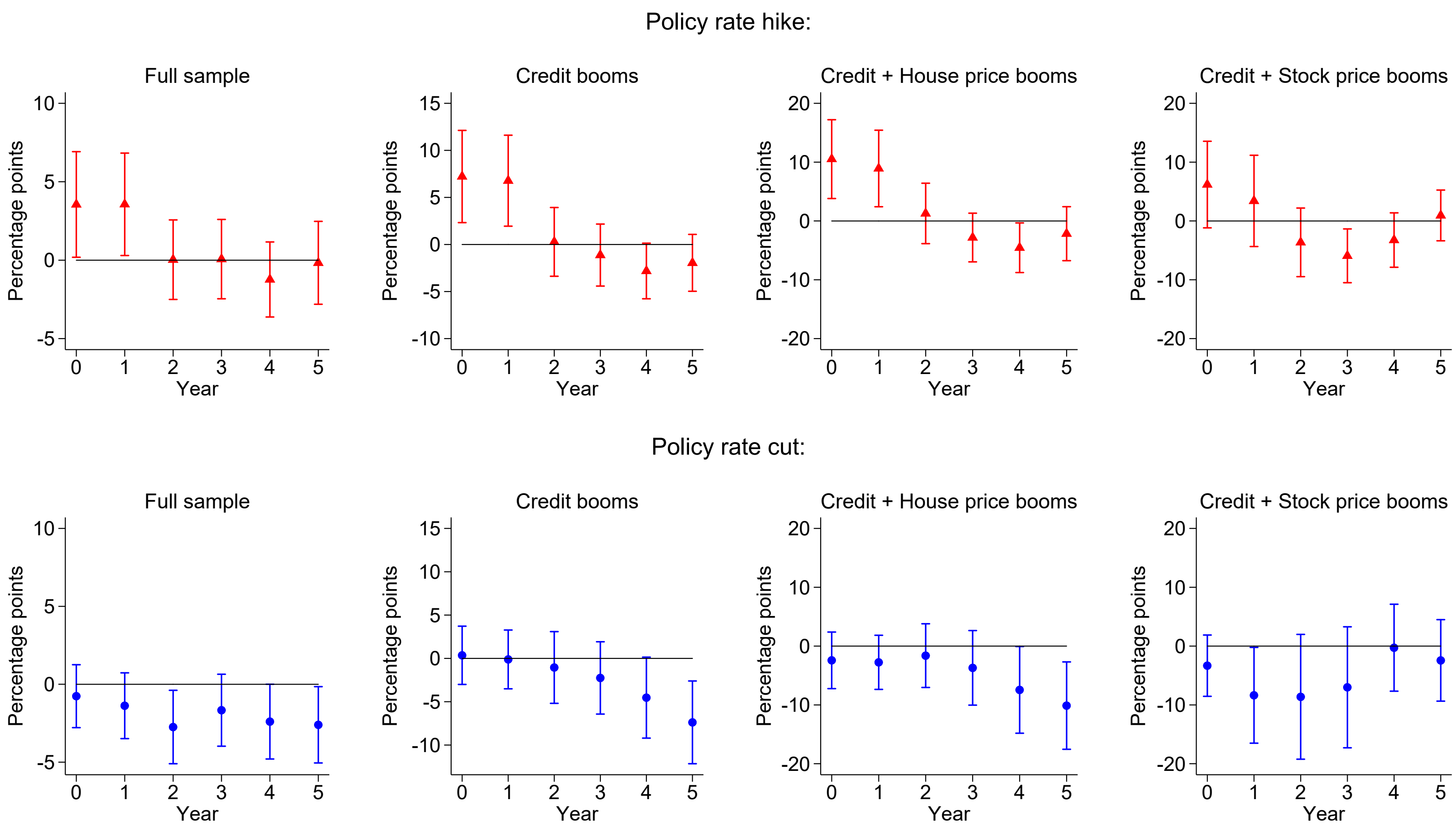
Notes: 95% confidence bands. Boom episodes defined on the basis of CF-filtered series (2- to 32-period range).

Figure 2.A.4: Crisis risk response—Hamilton filter



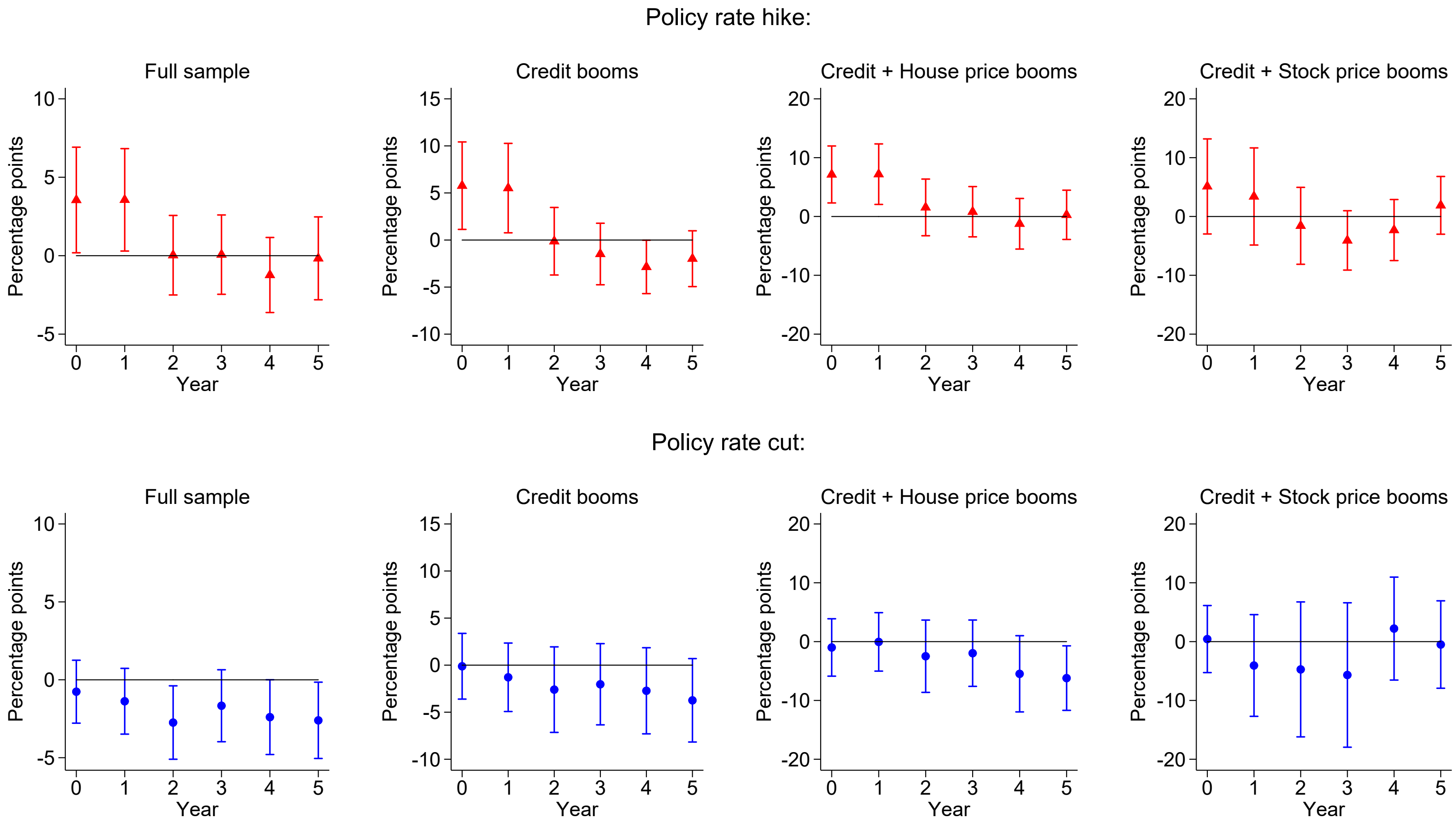
Notes: 95% confidence bands. Boom episodes defined on the basis of Hamilton-filtered series (lags 3 to 6).

Figure 2.A.5: Rate hikes versus rate cuts and crisis risk—Christiano-Fitzgerald filter



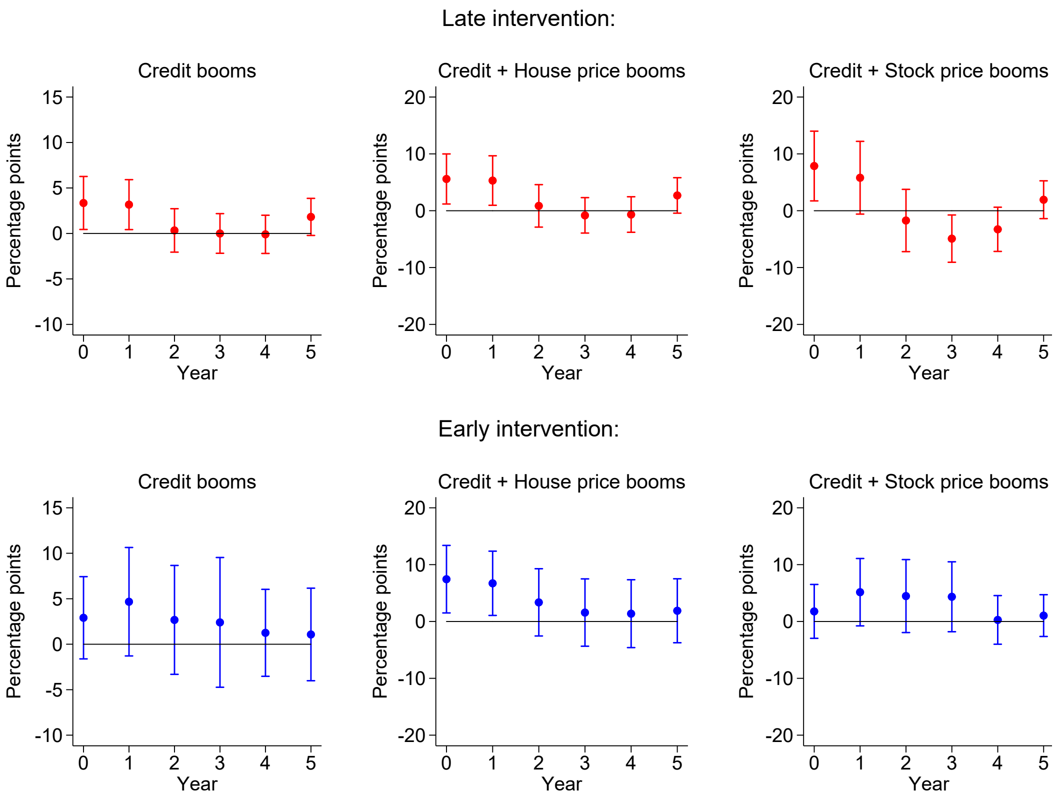
Notes: Change in the annual crisis probability following a 1 ppt policy rate hike/cut. 95% confidence bands.

Figure 2.A.6: Rate hikes versus rate cuts and crisis risk—Hamilton filter



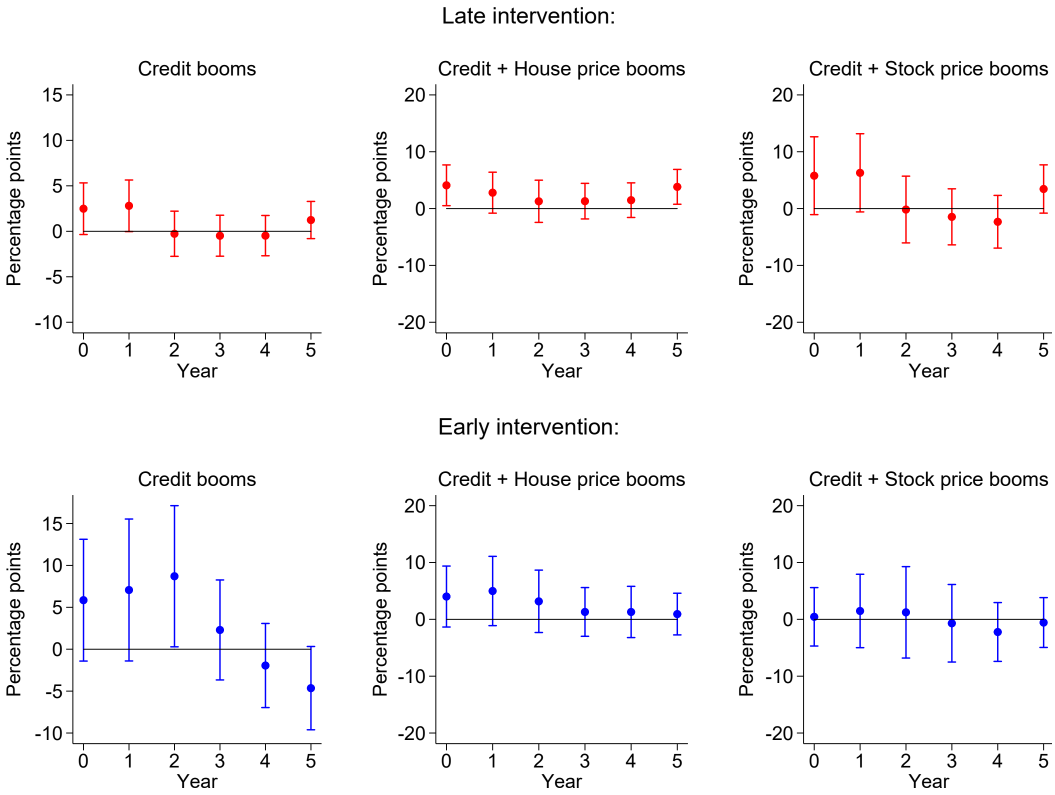
Notes: Change in the annual crisis probability following a 1 ppt policy rate hike/cut. 95% confidence bands.

Figure 2.A.7: Early versus late interventions and crisis risk—Christiano-Fitzgerald filter



Notes: Change in the annual crisis probability following a 1 ppt policy rate hike. 95% confidence bands.

Figure 2.A.8: Early versus late interventions and crisis risk—Hamilton filter



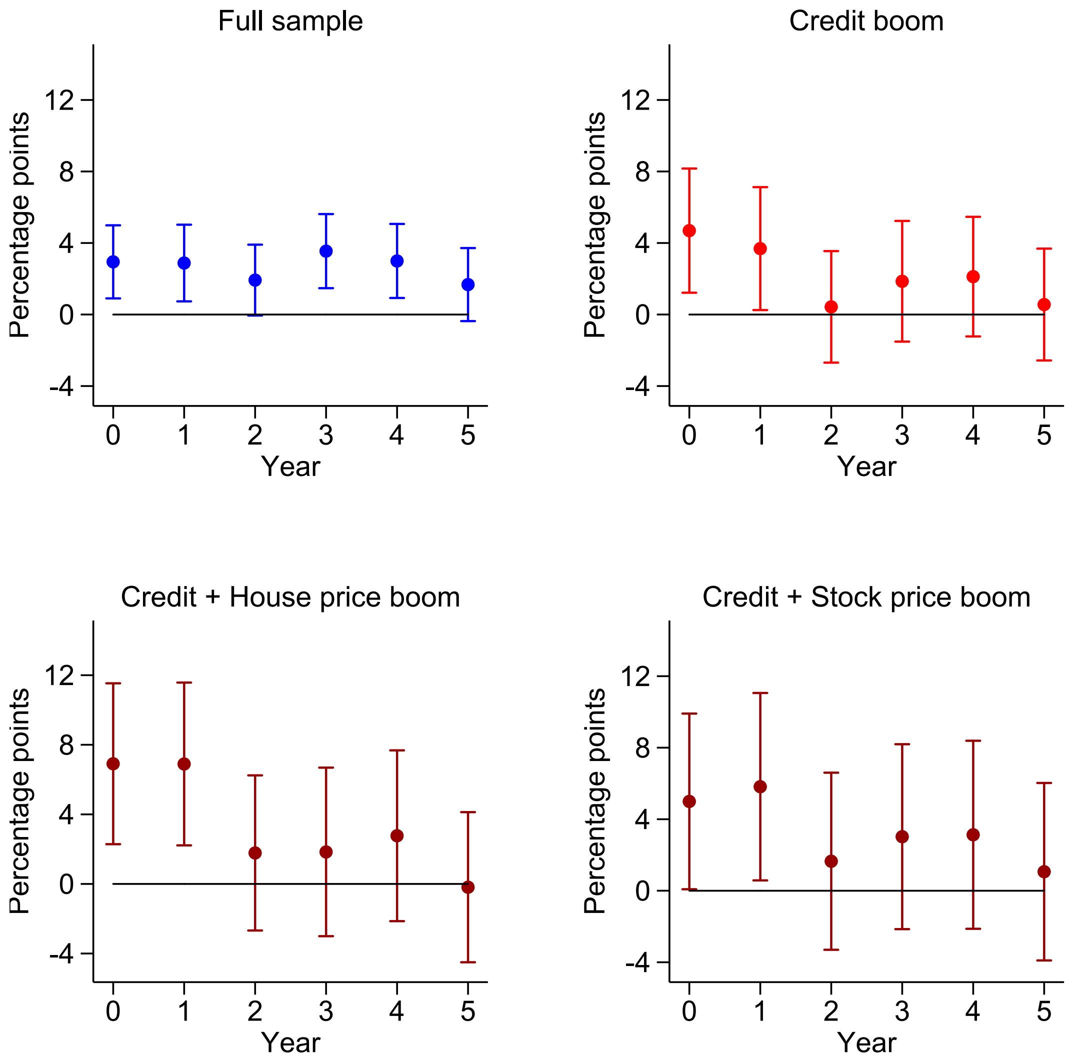
Notes: Change in the annual crisis probability following a 1 ppt policy rate hike. 95% confidence bands.

2.B Alternative financial crisis indicators

The way to date financial crises is always subject to debate, and different indicators of financial crises have been proposed in the literature. Our main results use the systemic financial crisis dummy from Schularick and Taylor (2012). One possible alternative to this indicator is the banking crisis dummy defined by Reinhart and Rogoff (2011). Their indicator marks a given year in a country as a banking crisis if either bank runs occurred that lead to closure, merging, or takeover by the public sector of one or more financial institutions, or, absent bank runs, the closure, merging, takeover, or large-scale government assistance of one or more important financial institutions marks the beginning of similar outcomes for other financial institutions.

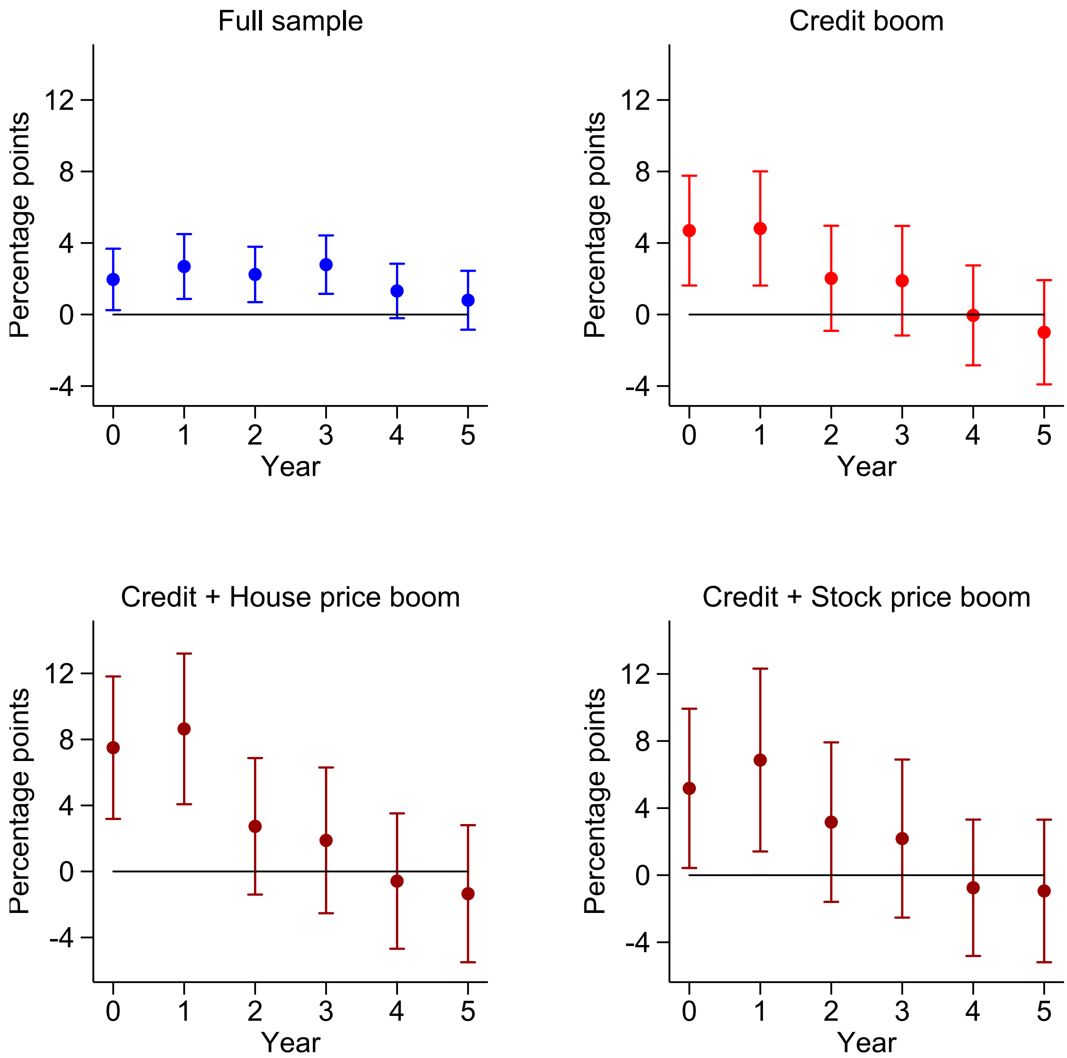
As a second robustness check, we also estimate the models using the dummy defined by Baron et al. (2018). Their approach to date banking crises creates a joint list of crisis dates from several studies and refines this list using data on bank equity declines. More precisely, the joint list is refined by adding episodes in which both the bank equity index declines by 30% or more and the narrative record shows substantial evidence of widespread banking failures or bank runs. These two criteria are also used to remove crisis dates from the joint list if neither condition is met. Thus, in contrast to the crisis dummy used in the main text as well as the crisis dummy defined by Reinhart and Rogoff (2011), this approach goes beyond the narrative identification by adding additional quantitative requirements to date financial crises. Figures 2.B.1-2.B.4 show that both the crisis risk responses and the crisis severity results from the main text are robust to using either indicator.

Figure 2.B.1: Crisis risk response—Reinhart and Rogoff crisis dummies



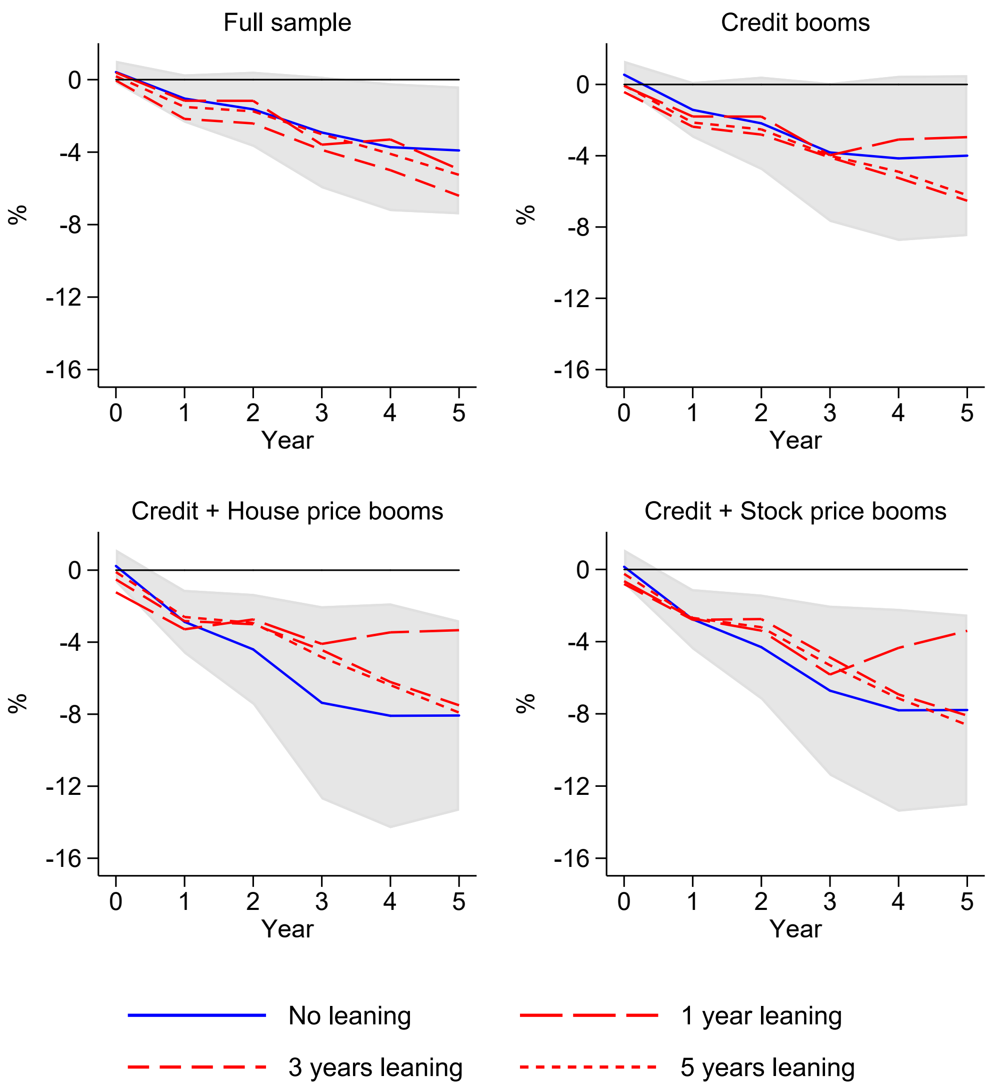
Notes: 95% confidence bands. Banking crisis dummy from Reinhart and Rogoff (2011).

Figure 2.B.2: Crisis risk response—Baron, Verner & Xiong crisis dummies



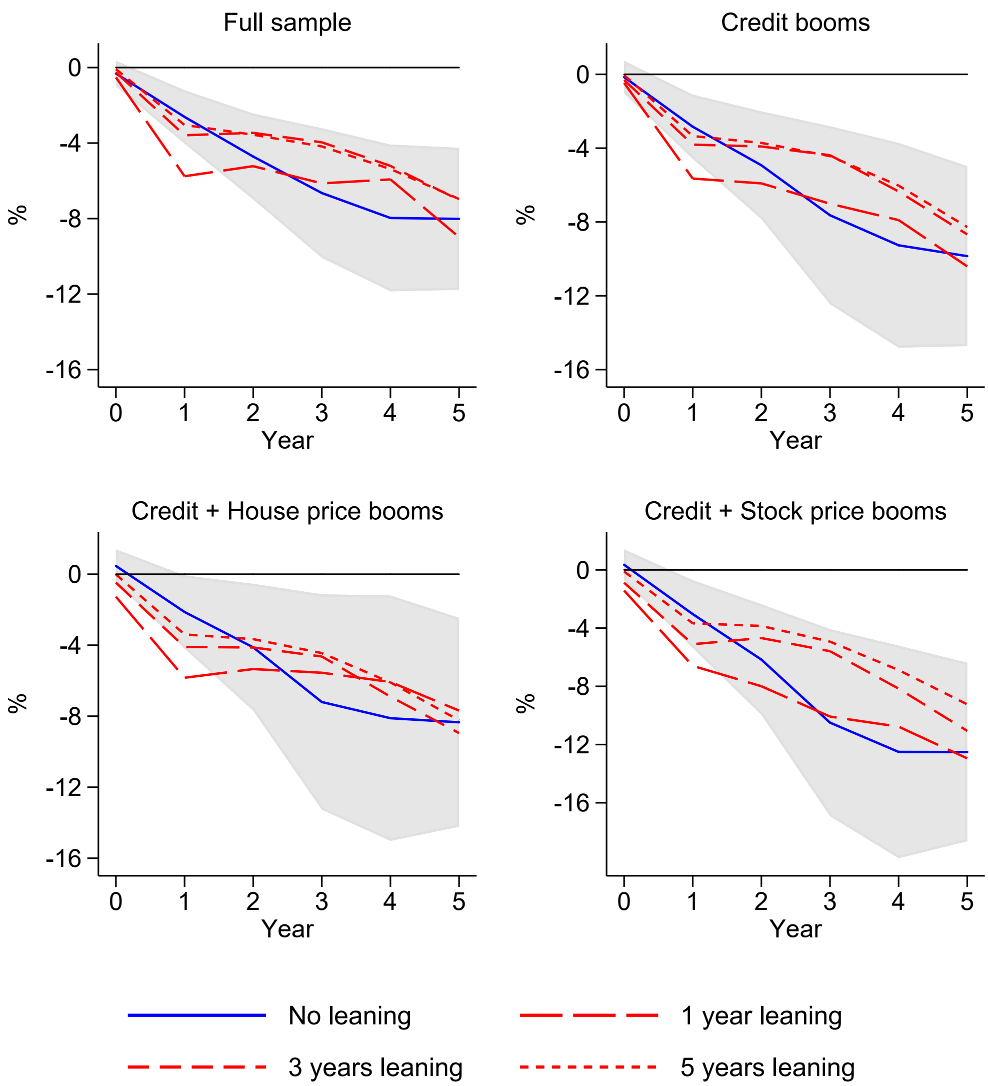
Notes: 95% confidence bands. Banking crisis dummy from Baron et al. (2018).

Figure 2.B.3: LAW and crisis severity—Reinhart and Rogoff crisis dummies



Notes: Real GDP loss after a crisis, depending on whether monetary policy was leaning against the wind or not. 95% confidence bands.

Figure 2.B.4: LAW and crisis severity—Baron, Verner & Xiong crisis dummies



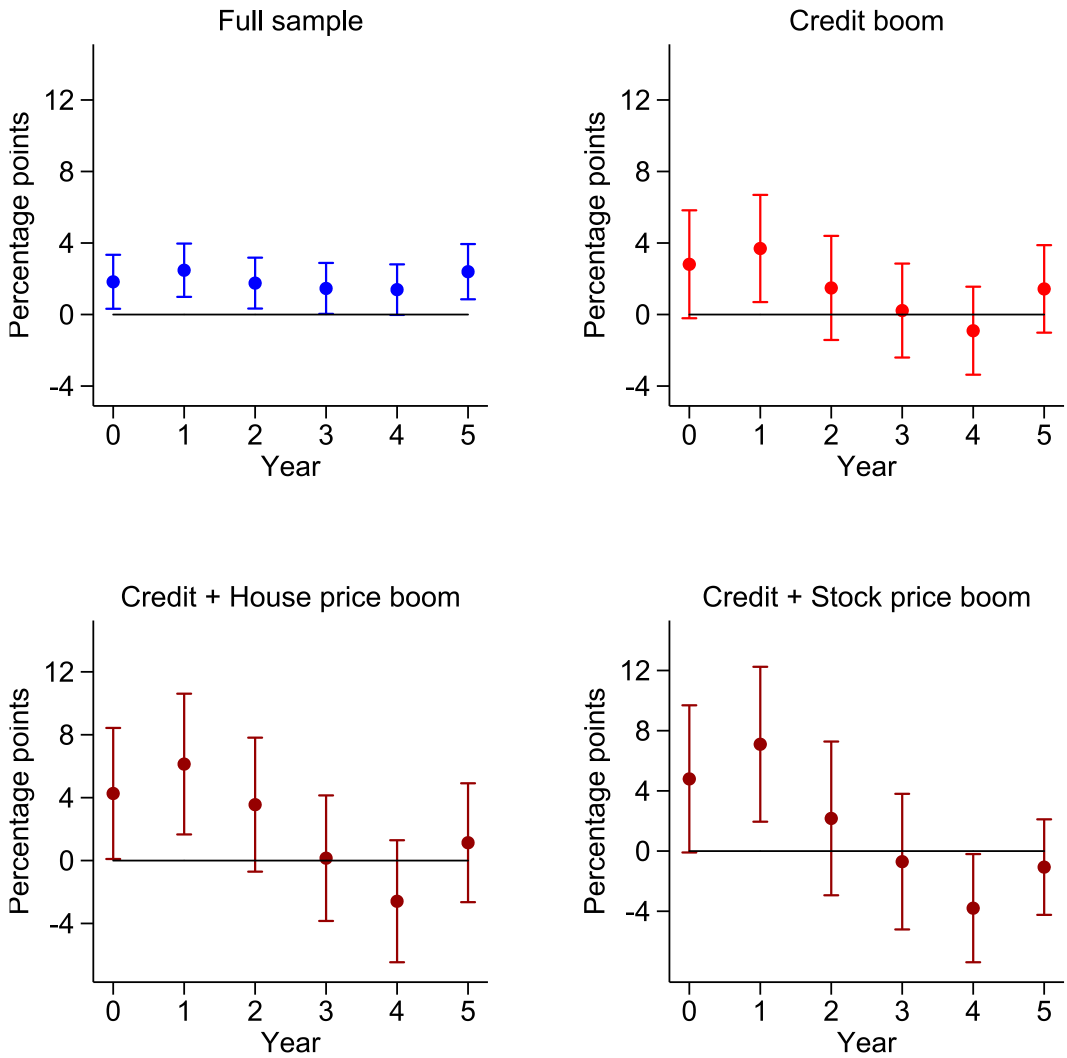
Notes: Real GDP loss after a crisis, depending on whether monetary policy was leaning against the wind or not. 95% confidence bands.

2.C Subsample analysis and model specification

Our final robustness checks are concerned with the post-World War II subsample, the model specification, the definition of early and late intervention in section 2.4.1, and potential spillover effects. First, Figure 2.C.1 shows that our main result regarding the crisis risk response is robust to restricting the sample period. Although somewhat smaller in size, the unconditional and boom episode responses show the same patterns as in the main text. Second, Figure 2.C.2 shows that our baseline conclusions remain valid when we switch from a linear probability model to a logit specification. Third, regarding early versus late interventions, the main text defines an intervention as early if it occurs in the first year of a boom period. Figure 2.C.3 and Figure 2.C.4 show, respectively, that our main results are robust to changing this window to include the first two years of a boom phase, or to include the first half.

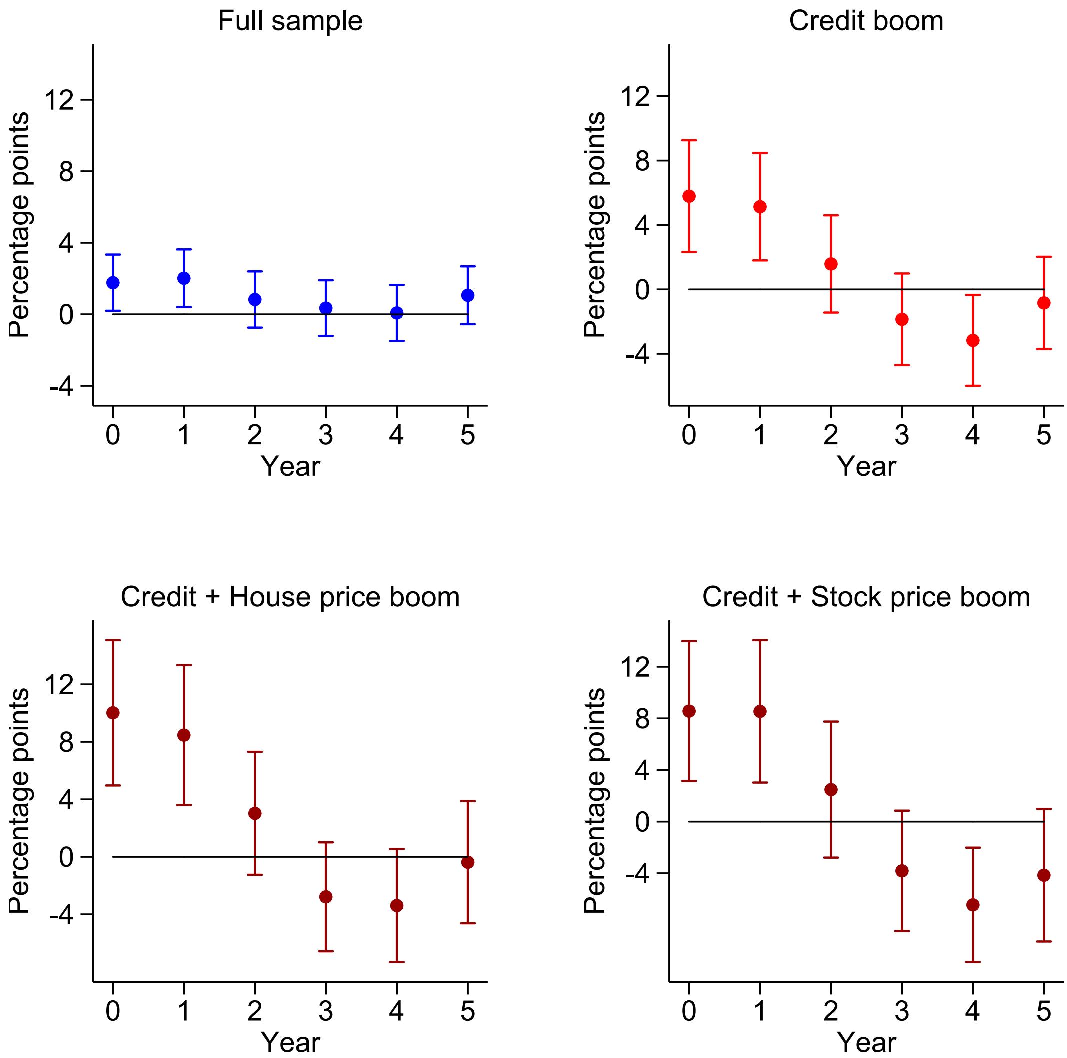
Last, we run the exercise from Jordà et al. (2019) to check whether potential spillover effects drive our results. For details we refer to their paper, but the intuition is as follows. The instrumental variable approach assumes that the instrument affects the outcome only through its effect on the policy rate. This assumption may plausibly be violated because of trade and financial market linkages between the countries in our sample. The idea is then to estimate this spillover effect from the sample of floating exchange rate countries, and with this estimate correct the IV estimates. Figure 2.C.5 shows that our baseline results are not driven by any potential bias arising from failure of the exclusion restriction.

Figure 2.C.1: Crisis risk response—Post-World War II subsample



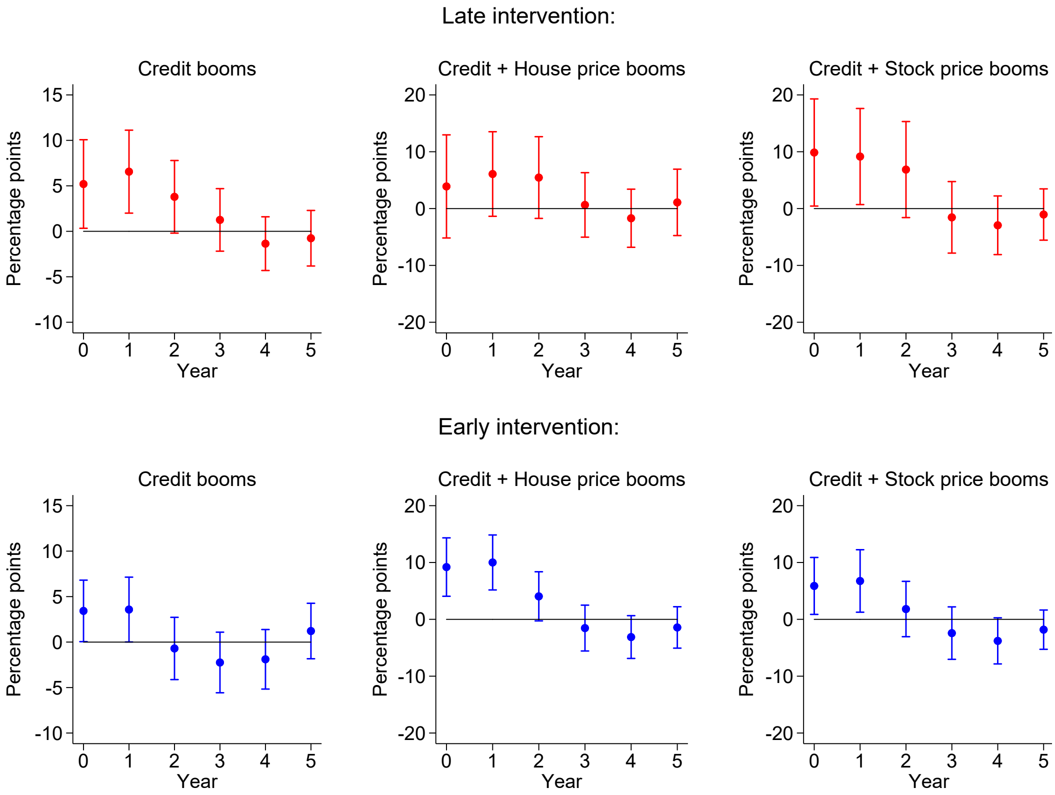
Notes: 95% confidence bands. Sample period restricted to the period after World War II. Boom episodes defined on the basis of HP-filtered series (λ=100).

Figure 2.C.2: Crisis risk response—Logit model



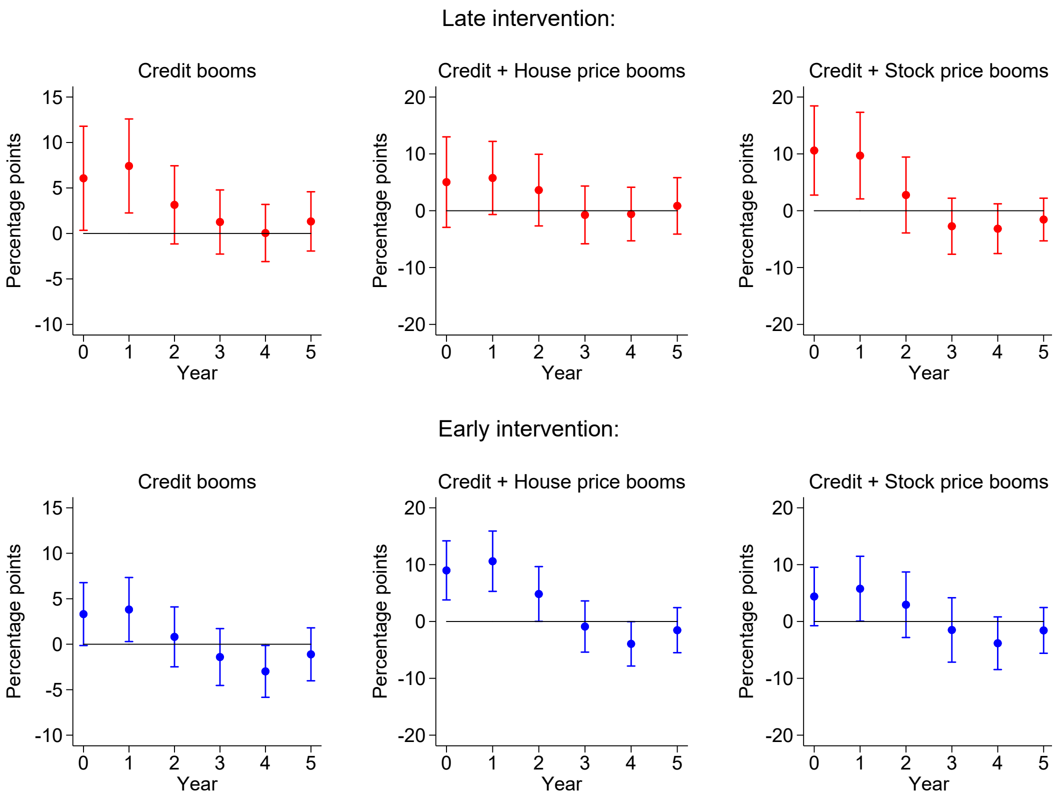
Notes: 95% confidence bands. Boom episodes defined on the basis of HP-filtered series (λ=100).

Figure 2.C.3: Early versus late interventions and crisis risk—First two boom years



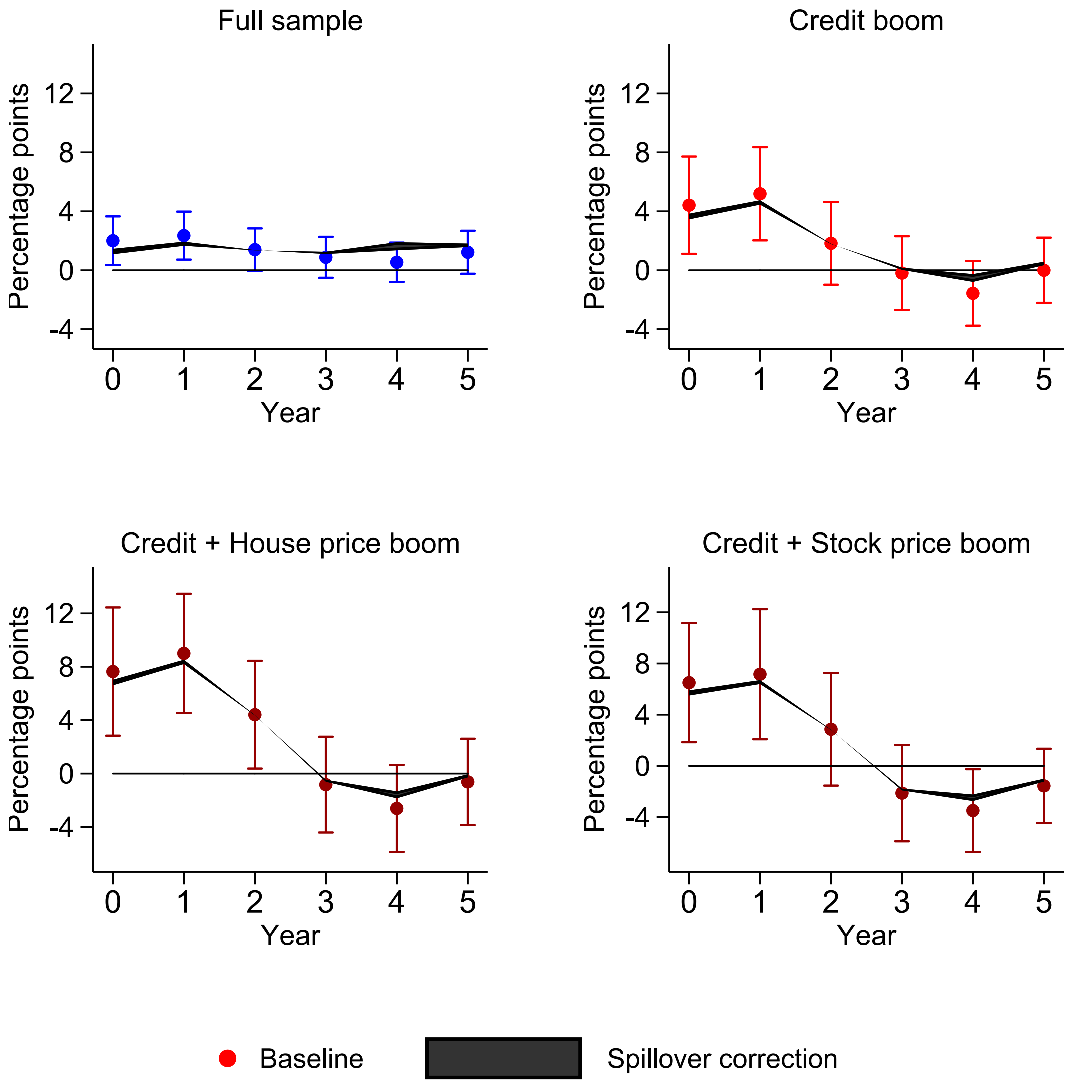
Notes: Change in the annual crisis probability following a 1 ppt policy rate hike. 95% confidence bands.

Figure 2.C.4: Early versus late interventions and crisis risk—First versus second boom halves



Notes: Change in the annual crisis probability following a 1 ppt policy rate hike. 95% confidence bands.

Figure 2.C.5: Crisis risk response—Spillover correction



Notes: Black area indicates range for spillover corrected mean IRF estimates. 95% confidence bands. Boom episodes defined on the basis of HP-filtered series (λ=100).

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Chapter 3

Destabilizing Effects of Consumer Bankruptcy [*]

Joint with Lisa Dähne

3.1 Introduction

Credit conditions are essential determinants in household saving and consumption decisions and are important drivers of business cycle fluctuations. The cost of and access to credit shapes the ability of households to smooth out adverse income shocks over time, and thus determines how individuals and the economy as a whole respond to macroeconomic shocks. Especially for low-income households who are likely to be financially constrained, a reduction in credit access has detrimental effects, as they see their ability to smooth consumption severely hampered. These are households who mostly rely on labor income the risk of which fluctuates largely over the business cycle and rises substantially during recessions (Bayer et al., 2019; Guvenen et al., 2014; Storesletten et al., 2004). In this paper, we ask whether these fluctuations in risk through effects on the price of unsecured credit have strong business cycle consequences. After an increase in risk, the likelihood of default in the future increases. Banks reflect this risk ex-ante and effectively limit the menu of available debt contracts by raising interest rates. This forces a deleveraging process on the part of borrowers through either declaring bankruptcy or a cutback in consumption. Both lead to a reduction of assets in the economy that can accommodate savings, and thus conversely to an excess supply of goods. If output is demand determined, a recession impulse follows.

Specifically, we first show empirically that borrowing conditions in the credit card sector comove strongly with income risk over the business cycle and that these

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conditions become tighter in response to income risk shocks. During major recessions, banks tighten their lending standards very early on, at the same time that income risk increases. This lends support for the hypothesis that banks preemptively tighten credit constraints in anticipation of future charge-offs, which indeed rise dramatically over the course of recessions. We further estimate the effects of income risk shocks using Local Projections as in Jordà (2005). In line with the descriptive evidence, we find that lending standards tighten on impact, while credit card charge-offs respond with a delay. Additionally, consumer credit becomes more costly relative to savings and consumer credit declines, consistent with the idea that tighter borrowing constraints force household to adjust to lower debt levels.

These results are then explored formally in a structural model of consumer default embedded in a Heterogeneous Agent New Keynesian (HANK) framework. This model captures the traditional precautionary savings channel through which households increase their savings in response to an increase in income risk, yet allows us to also assess how such risk affects borrowing limits and amplifies household behavior. Consumer default is modeled along the lines of Chatterjee et al. (2007). In the model, competitive banks price individual loans according to the likelihood of default based on an individuals observable characteristics, in line with empirical evidence in Edelberg (2006). As income risk increases, so does the default probability, which banks compensate by charging higher interest rates on loans, thus making borrowing more expensive.

Our analysis shows that this channel is quantitatively important and generates larger consumption and output drops compared to standard HANK models in which consumer bankruptcy does not exist. This is because the endogenous tightening of borrowing constraints adds to precautionary savings the fact that households who want to borrow also have to reduce their debt position, yet do not find it optimal to default. We find that the endogenous response of the price of borrowing accounts for about a third of the overall drop in output and consumption. As a result, rather than increasing the ability for households to smooth consumption, consumer bankruptcy makes the cost of borrowing and hence aggregate consumption more volatile in response to income risk shocks.

Our findings complement recent papers that study ex-post insurance provided by unsecured credit. Auclert et al. (2019) study the role of consumer debt relief during the Great Recession and document a less pronounced drop in non-traded employment, higher write-offs of unsecured debt in U.S. states with more generous bankruptcy exemptions, and an increase in consumption and employment in response to debt relief. They model debt relief as a sequence of transfers from savers to borrowers that is specifically designed to not affect ex-ante credit supply and the cost of borrowing. In contrast, Herkenhoff (2019) shows that access to credit can lead to protracted recessions as job search by the unemployed increases. Borrowers who are hit by unemployment spells choose default to smooth consumption, and access to credit offers a safety net that job searchers take advantage of to increase time spend looking for a job. Our paper instead studies the ex-ante effects of news about increased income risk in the future on the cost of borrowing, and how this effects the business cycle.

The way monetary policy reacts to inflation plays an important role in our environment. By cutting interest rates the central bank leans against the risk-induced increase in borrowing rates. We find that a more aggressive monetary policy response expands access to credit for very poor households, and significantly reduces overall tightening of borrowing constraints, which directly translates into reduced default rates and reduced delevering in the economy. The result that low-income households see their credit limits expand comes from the fact that these are high-risk households with little access to credit to begin with, consistent with evidence in Sullivan (2008). Credit limits for these households are less sensitive to changes in income risk and more sensitive to variations in the central bank interest rate. However, we also find that when the central bank sharply cuts interest rates on debt, it later has to raise interest rates above their steady state values so as to later on reign in the initial expansion in household debt that comes with the sharp initial cut in interest rates.

The mechanism we highlight is similar to Guerrieri and Lorenzoni (2017) who study the effects of an exogenous credit crunch in a heterogeneous agent incomplete market economy and find that the resulting deleveraging process strongly depresses economic activity. We generate these effects in a model of consumer bankruptcy, which endogenously generates a credit crunch due to an increase in household income risk. Our paper is also closely related to the literature studying consumer default and business cycles. Nakajima and Ríos-Rull (2014) in a real business cycle model with counter-cyclical earnings risk find that during recessions, the default premium and hence borrowing rates rise, deterring household borrowing, which makes consumption more volatile. A similar result that unsecured credit markets do not provide better insurance opportunities is made in Athreya et al. (2009). Fieldhouse et al. (2016) investigate whether the framework of Livshits et al. (2007) when augmented by aggregate risk can replicate cyclical patterns of consumer credit and bankruptcy. Gordon (2015) studies the welfare effects of different bankruptcy regimes when the economy is subject to aggregate risk. Our paper is complementary to the existing literature in that we integrate consumer default into a New Keynesian model where output is demand determined. This allows us to investigate how higher income risk generates large recessions due to reductions in aggregate demand following endogenously tightening borrowing constraints, and in particular, how these effects interact with monetary policy.

The remainder of this paper is organized as follows. In section 3.2 we present empirical evidence on the comovement of household income risk and credit conditions and estimate the effect of income risk shocks on economic activity and credit conditions. Section 3.3 describes the model environment, the solution method, and the calibration. Section 3.4 presents the numerical results under our baseline calibration. In section 3.5 we study role of stabilization policy by the central bank as well as the transmission of monetary policy shocks. Section 3.6 concludes.

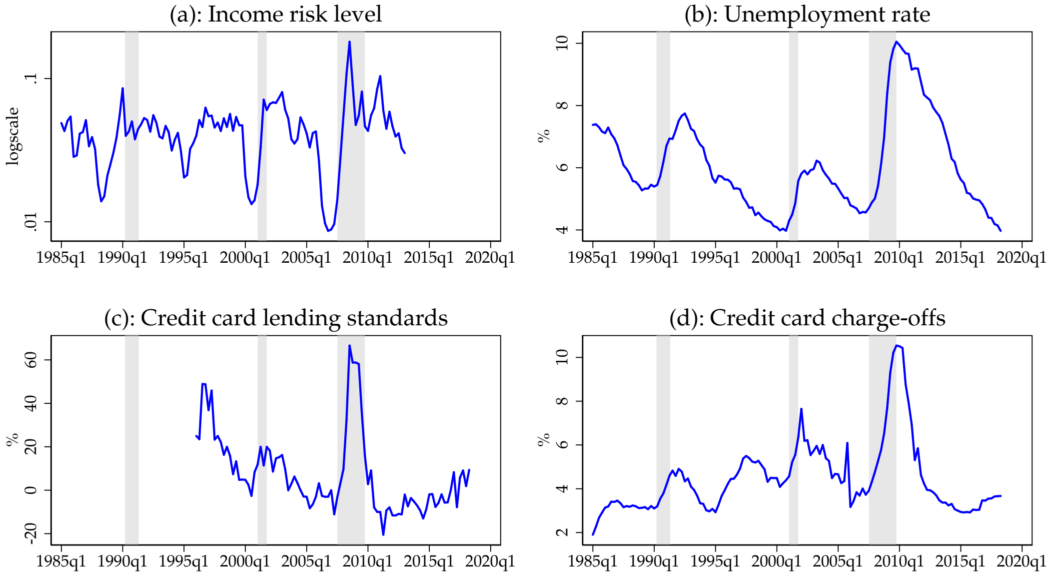
3.2 Empirical evidence

In this section, we investigate the aggregate dynamics of household income risk and credit conditions, and the responses of credit conditions to income risk shocks. Since our quantitative model is one of unsecured credit, we focus here on credit card loans by commercial banks. As our empirical measure of household income risk, we use the estimates from Bayer et al. (2019). They use household level data on income after taxes and transfers from the Survey of Income and Program Participation to estimate a rich income process that includes a time varying variance of shocks to the persistent component of income. This provides us with time series for the variance of household income as well as shocks to this variance, which we call income risk shocks. In addition to income risk, we use aggregate time series from the FRED database on total consumer credit, credit card charge-off rates, lending standards, and interest rates on credit card plans. Lastly, we also use time series for the civilian unemployment rate, real GDP, real consumption, real investment, the GDP deflator, and the effective federal funds rate. All time series are of quarterly frequency. Our sample covers the period from the first quarter of 1985 to the second quarter of 2018.

To start, in panels (a) and (b) of Figure 3.1 we plot the variance of the persistent component of household income together with the unemployment rate. Income risk is clearly counter-cyclical as is of course the unemployment rate. The variance of persistent income substantially rises at the start of recessions, while unemployment increases with a lag. Panels (c) and (d) depict credit card lending standards and charge-off rates on credit card loans by commercial banks. The former is defined as the net percentage of commercial banks tightening their lending standards on credit card loans. Clearly, both lending standards and charge-off rates rise during recessions, as banks see the repayment capabilities of their borrowers diminish and consumers either cannot repay their debts or find it optimal not to do so[1]. What is interesting here is that lending standards increase early during recessions, while

[1]The spike visible in charge-off rates in 2005 is due to the enactment of the Bankruptcy Abuse Prevention and Consumer Protection Act (BAPCPA), which made it considerably more difficult to file for bankruptcy, and many consumers rushed to file before the law was passed (Nakajima, 2017).

Figure 3.1: Income risk and credit card lending



Notes: Income risk level denotes the standard deviation of persistent income shocks from Bayer et al. (2019). The remaining series come the FRED database. Shaded areas denote NBER recessions.

charge-off rates show a build-up over the course of recessions, similar to the unemployment rate. This indicates that banks anticipate future financial distress of their borrowers and tighten credit standards preemptively. By contrast, consumers seem to take advantage of the option to discharge debt only at later stages, once the drop in income and employment, coupled together with limited borrowing opportunities, has made declaring bankruptcy either optimal or inevitable. The preemptive tightening of lending standards is also at the heart of models of consumer bankruptcy such as ours in which financial intermediaries price unsecured loans based on future repayment probabilities.

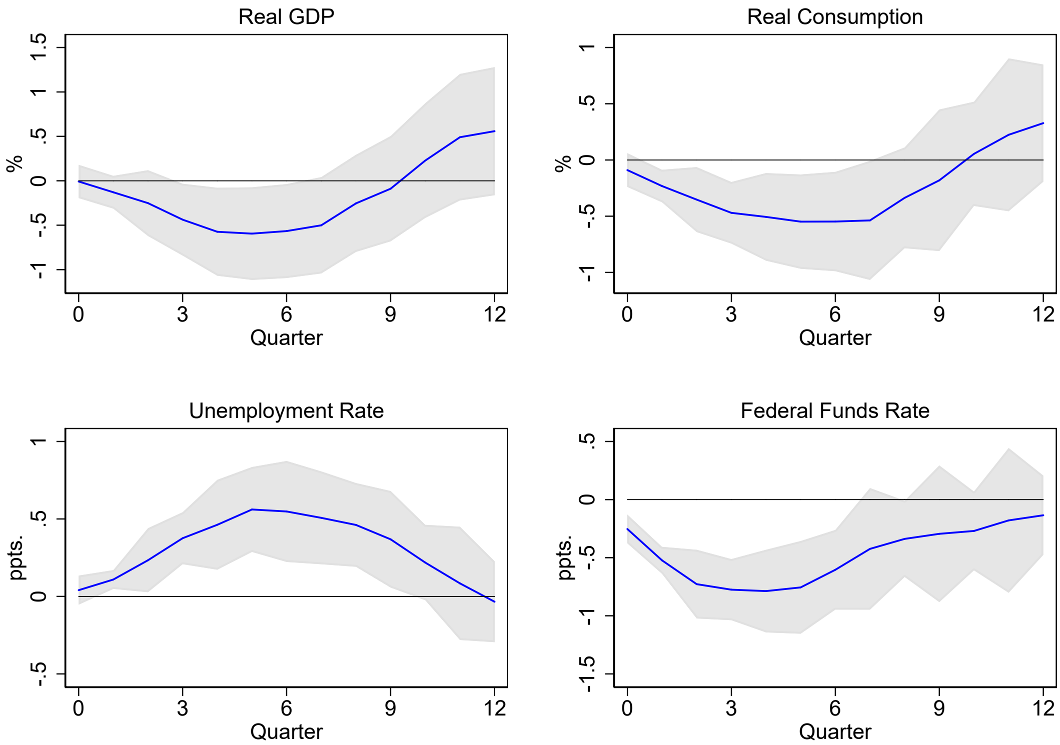
Next, we turn to the impact of income risk shocks on macroeconomic variables. To this end, we estimate impulse response functions with Local Projections (Jordà, 2005) using the income risk shocks series from Bayer et al. (2019) as our impulse variable. More specifically, we estimating the following set of regressions for each horizon $h = 0, \ldots, H$:

$$y_{t+h} = \alpha_h + \theta_h \epsilon_t + \Gamma_h y_{t-1} + u_{t+h} \tag{3.1}$$

y_{t+h} denotes the dependent variable of interest h steps ahead, and ϵ_t denotes the income risk shock. Equation (3.1) also includes the past value of the dependent variable to control for potential nonstationarity of the respective variable[2]. The

[2]This specification includes y_{t-1} only to control for nonstationarity. We have also estimated the

Figure 3.2: Responses to income risk shocks



Notes: Income risk shocks come from Bayer et al. (2019), the remaining series come the FRED database. Shaded areas denote 90% error bands constructed using Newey-West standard errors.

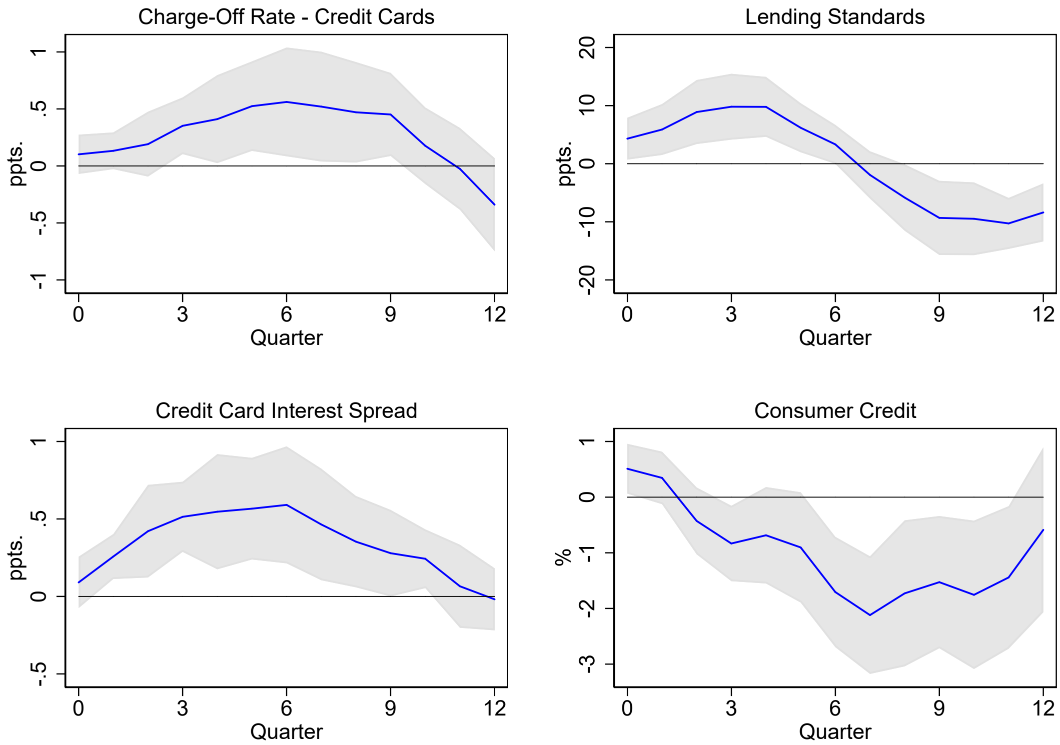
effect of income risk shocks on the outcome variable are then directly given by the coefficients θ_h. To improve the efficiency of the point estimates, for $h > 1$ we include the residual from the previous regression as an additional regressor (Jordà, 2005). In the following, all variables except interest rates, the unemployment rate, and the net percentage of banks increasing lending standards, are transformed into 100 times the natural logarithm of that variable.

Figure 3.2 largely replicate the empirical evidence from Bayer et al. (2019) and shows that both real GDP and real consumption fall within the first year and reach their trough around 4 quarters after the shock. GDP falls by about 0.5% at the trough, and the consumption response is slightly smaller. Similarly, the unemployment rate rises by 0.5 percentage points, while the nominal interest rate declines by about 0.8 percentage points within the first year, before returning to the pre-shock level.

Figure 3.3 adds to this and plots responses of variables related to availability and the cost of consumer credit: the charge-off rates and lending standards from Figure

model in cumulated first differences without controls as well as with a richer control set, with very similar results. These can be found in Appendix 3.B.

Figure 3.3: Responses of financial variables to income risk shocks



Notes: Income risk shocks come from Bayer et al. (2019), the remaining series come the FRED database. Shaded areas denote 90 % error bands constructed using Newey-West standard errors.

3.1 as well as the difference between interest rates charged on credit card plans and the federal funds rate, intended to capture premia on unsecured credit, as well as total liabilities from consumer credit of households and nonprofit organizations. Evidently, following income risk shocks, we find an increase in credit card default rates and tighter lending conditions. In line with the evidence in Figure 3.1, we observe that lending standards jump upward on impact and increase to their peak much faster than charge-off rates do, while charge-off rates build up slowly over time. Additionally, while the interest rate charged on credit card plans falls (not shown), the nominal central bank interest rate declines faster, leading to an increased spread between the two. Lastly, there is a prolonged reduction in consumer credit, which falls by 2% after seven quarters before slowly recovering to trend.

Taken together, the patterns established point to the following mechanisms as a possible explanation. Banks rationally price the increased income and default risk and retract lending to higher risk households, thus leading to an increased spread between risky and safe borrowing rates as well as tighter lending standards. At the same time, households also cut back on consumption so as to increase their precautionary savings to accumulate more buffer stock savings against future income shocks. The decline in aggregate demand depresses production and employment, to which the central bank endogenously responds by cutting interest rates to offset some of the economic decline. Low income households may then find it optimal or necessary to not repay their debts, increasing default rates in the economy. The endogenous increase in savings, higher credit costs, and higher default rates then leads to the observed decline in consumer credit. In the following we explore these mechanisms in a quantitative model that incorporates unsecured credit, time-varying household income risk, and sticky prices.

3.3 Model

To study how time-varying income risk effects the price of unsecured credit and how this contributes to business cycle fluctuations, we build a model with heterogeneous agents, incomplete markets, consumer default, and sticky prices. Our model economy is composed of a household sector, a firm sector, a banking sector, and a government sector. Households supply labor to firms, save in one-period defaultable bonds, and own the final-goods firms and banks. Firms either differentiate intermediate goods into final goods facing monopolistic competition and pricing frictions, or produce intermediate goods under perfect competition. Banks issue loans to borrowers and price these loans according to borrower characteristics in a perfectly competitive market. The fiscal side of the government levies an income tax, issues bonds, and adjusts expenditures in reaction to changes in government debt and business cycle conditions. The monetary authority controls the nominal interest rate according to a Taylor rule.

3.3.1 Preferences and productivity process

Households are ex-ante identical, infinitely lived, and have time-separable preferences over consumption and leisure services. We assume that the period utility function over both consumption and leisure services features constant relative risk aversion

$$u(x_{it}) = \frac{1}{1-\xi} x_{it}^{1-\xi} \tag{3.2}$$

$$x_{it} = c_{it} - G(h_{it}, n_{it}) \tag{3.3}$$

$$G(h_{it}, n_{it}) = h_{it} \frac{n_{it}^{1+\gamma}}{1+\gamma} \tag{3.4}$$

where c_{it} is real consumption of household i at time t, $G(h_{it}, n_{it})$ denotes disutility from labor, and $\xi > 0$ and $\gamma > 0$ denote the relative risk aversion and inverse Frisch elasticity parameters, respectively. Lastly, h_{it} denotes individual labor productivity.

Parameterizing the disutility of labor in hours worked and labor productivity implies that individual labor supply depends on the aggregate wage rate only, that is $n_{it} = N(w_t)$. It also implies that we can write the budget constraint in terms of aggregate labor and combined consumption and leisure services, x_{it}.

Regarding individual labor productivity, we adopt the productivity process from Bayer et al. (2019). Households can either be workers or entrepreneurs. When working, individual labor productivity evolves according to an AR(1) process with normally distributed innovations. The variance of these innovations is time-varying. Entrepreneurs are unproductive in the workforce and therefore do not receive income from working. Instead, entrepreneurs receive firm profits that arise in the final goods sector and profits that arise in the banking sector. With small probability, workers become entrepreneurs, and entrepreneurs that become workers again enter the workforce with average productivity. The law of motion for productivity is

$$\tilde{h}_{it} = \begin{cases} \exp(\rho_h \log(\tilde{h}_{it-1}) + \epsilon^h_{it}) & \text{with probability } 1-\zeta \text{ if } h_{it-1} \neq 0, \\ 1 & \text{with probability } \iota \text{ if } h_{it-1} = 0, \\ 0 & \text{otherwise.} \end{cases} \tag{3.5}$$

where individual productivity is $h_{it} = \frac{\tilde{h}_{it}}{\int \tilde{h}_{it} di}$ so that average productivity is constant. The variance of the shocks ϵ^h_{it} evolves according to

$$\sigma^2_{ht} = \bar{\sigma}^2_h \exp(S_t) \tag{3.6}$$

$$S_{t+1} = \rho_S S_t + \epsilon^S_t \tag{3.7}$$

$$\epsilon^S_t \sim \mathcal{N}\left(-\frac{\sigma^2_S}{2(1+\rho_S)}, \sigma^2_S\right) \tag{3.8}$$

Thus, aggregate shocks to future household productivity are observed at time t which can also be interpreted as news about the future variance of individual productivity. These shocks are also observed by banks, who use this information to correctly price individual default risk, as we show in the next section.

3.3.2 Bankruptcy and financial markets

Following Chatterjee et al. (2007) and many others, we model default as being Chapter 7 default only. The data from the United State Court system show that within each year, about 60-70 % of all consumer bankruptcy cases are filed under Chapter 7, so we think this is a reasonable simplification of the problem. Upon declaring bankruptcy, a household's debt is wiped out, and no financial market transaction is allowed during the period of default. In addition, the household has to pay a utility cost and receives a default flag. With a default flag, a household enters a state in which it is prohibited from borrowing. With a constant probability the household leaves the default state every period, the default flag is erased, and full participation in financial markets is possible again.

In the model there exists a perfectly competitive banking sector that takes deposits from savers and makes loans to borrowers. In case of default, banks receive nothing. Banks observe an individual's productivity h_{it} and know the productivity process. They will therefore use this information to infer the probability that the loan is not repayed in the next period. A loan is a promise to repay an amount $b_{it+1} < 0$ next period, for which the borrower will receive $q_t(b_{it+1}, h_{it})b_{it+1}$ this period. Perfect competition and the default option imply that the price for a loan of size b_{it+1}, given individual productivity h_{it}, is

$$q_t(b_{it+1}, h_{it}; S_t) = \frac{1}{R_{t+1}} \left[1 - \mathbb{P}(d_{it+1} = 1 | b_{it+1}, h_{it}; S_t)\right] \tag{3.9}$$

where $d_{it+1} = 1$ indicates default in the next period, and R_{t+1} is the nominal gross interest rate set by the central bank. Accordingly, banks will offer the risk-free nominal bond price $1/R_{t+1}$ whenever a household will never default next period, and charge a premium otherwise. The interest rate used to discount future repayments is set by the central bank. Savings by the government, denoted B_t, is priced the same way, although we abstract from government default, so that the default probability is always zero[3]. Also, banks observe the current aggregate risk state S_t and use it together with individual characteristics to infer the likelihood of default next period.

The bond pricing equation (3.9) reveals the key mechanism at play in our model. It is well known from models of sovereign default and consumer bankruptcy that the probability of default increases with debt (Arellano, 2008; Chatterjee et al., 2007; Mendoza and Yue, 2012). In our quantitative exercise, it is also the case that default incentives are stronger for lower productivity realizations. Suppose now that banks receive information that income risk increases, which increases the probability to receive a bad productivity realization next period. The bank will correctly anticipate that the default probability increases for a given level of debt and productivity realization, and hence the amount a borrower receives upon a given promise to repay declines. In contrast to this, if the central bank cuts interest rates, bond prices rise ceteris paribus, expanding the amount of resources receive for a given promise. Clearly, this effect applies uniformly to every loan contract, while the change in default probabilities is individual specific. Which of these effects dominates in general equilibrium depends on individual characteristics. As we will show below, the risk effect is stronger around the middle of the income distribution,

[3] We denote government savings by B_t, so that government debt is $-B_t$. We chose this notation to avoid switching signs when we speak of individual and government debt.

while in the bottom and high end the interest rate effect is stronger .

As with the price setting of final goods producers discussed in the next section, we assume that a mass zero households of risk neutral managers operates these banks, and all profits and losses in the financial sector go to the entrepreneur households. Real profits in the financial sector are given by

$$\Pi_t^b = -\left(\int \frac{b_{it}}{\pi_t} \mathbb{I}_{f_{it}=0} \mathbb{I}_{d_{it}=0} di + \int \frac{b_{it}}{\pi_t} \mathbb{I}_{f_{it}=1} di + \frac{B_t}{\pi_t} \right) \tag{3.10}$$

where $\pi_t = \frac{P_t}{P_{t-1}}$ denotes realized gross inflation. The first integral gives the total real amount that is paid out in period t to households that currently do not have a default flag ($\mathbb{I}_{f_{it}=0} = 1$) and do not default ($\mathbb{I}_{d_{it}=0} = 1$). The second integral gives the total real amount paid out to those that are borrowing constrained. Since the household sector is a net saver, the sum of these integrals is positive, and the government is a debtor, so that B_t is negative. In steady state without aggregate risk, the net amount paid to households is equal to repayment by the government, so that realized profits are always zero. With aggregate shocks, however, there can be a difference between the amount promised and actually repaid by households, and this difference is borne by the entrepreneurs.

3.3.3 Household planning problem

To spell out the household planning problem, define $\mathcal{Z}_t = (\Theta_t, R_t, S_t)$ as the collection of aggregate state variables in the economy. These are the joint distribution of households over wealth, productivity, and bankruptcy status, Θ_t, the central bank interest rate, R_t, and the aggregate uncertainty state S_t. Households then solve their maximization problems conditional on their idiosyncratic wealth b_{it} and productivity h_{it}, as well as $\mathcal{Z}_t$.

Let W^R denote the value function associated with repaying, W^D the value of defaulting, V the upper envelope between the two, and W^C the value obtained while in the borrowing constrained state. Dropping time indices and denoting next period values with a prime, the recursive planning problem then takes the form

$$\begin{aligned} W^R(b,h;\mathcal{Z}) &= \max_{x,b'} \{u(x) + \beta \mathbb{E}[V(b',h';\mathcal{Z}')]\} \\ s.t. \quad x + q(b',h;S)b' &= \frac{b}{\pi} + (1-\tau)\left(\frac{\gamma}{1+\gamma} whN(w) + \frac{1}{\Phi}\mathbb{I}_{h=0}\Pi\right) \end{aligned} \tag{3.11}$$

for households that choose to repay. In words, the problem in (3.11) states that a household that repays chooses real composite consumption x and real bonds b' at price $q(b',h)$ such that total real expenditures equal real income, which is composed of the real value of assets today and real after-tax labor and profit income. We define $b > 0$ as savings and $b < 0$ as liabilities, and $\mathbb{I}_{h=0}$ is an indicator whether the household is an entrepreneur or not. Profits are scaled by the mass of entrepreneur households, Φ, such that per-entrepreneur household profits enter in the individuals' budget constraint. The continuation value for this problem is the upper envelope of the repayment and default value $V(\cdot)$, which we define below, since the household can default tomorrow given repayment today.

The planning problem for defaulters takes the following form:

$$\begin{aligned} W^D(h;\mathcal{Z}) &= u(x) - \phi + \beta\mathbb{E}[W^C(0,h';\mathcal{Z}')] \\ s.t. \quad x &= (1-\tau)\left(\frac{\gamma}{1+\gamma}whN(w) + \frac{1}{\Phi}\mathbb{I}_{h=0}\Pi\right) \end{aligned} \tag{3.12}$$

Equation (3.12) states that upon defaulting, the asset position is wiped out, a default cost $\phi > 0$ is paid, the household consumes its income, and no asset market participation is possible in the period of defaulting. Modeling default costs this way is borrowed from Athreya (2002) and ϕ can be interpreted as stigma associated with declaring bankruptcy (see also Athreya (2004)). In the period following bankruptcy, the household starts with zero assets in the constrained state. In this state, the planning problem becomes

$$W^C(b,h;\mathcal{Z}) = \max_{x,b'}\{u(x) + \beta\mathbb{E}[\lambda V(b',h';\mathcal{Z}') + (1-\lambda)W^C(b',h';\mathcal{Z}')]\} \tag{3.13}$$

$$\begin{aligned} s.t. \quad x + \tilde{q}(b',h;S)b' &= \frac{b}{\pi} + (1-\tau)\left(\frac{\gamma}{1+\gamma}whN(w) + \frac{1}{\Phi}\mathbb{I}_{h=0}\Pi\right) \\ b' &\geq 0 \end{aligned}$$

In (3.13), households with a default flag choose between consumption and savings subject to a no-borrowing constrained. With probability λ, the household has the default flag removed next period and returns to problem (3.11). With the complementary probability $1-\lambda$, the household remains constrained next period.

Finally, the upper envelope of repaying and defaulting is given by

$$V(b,h;\mathcal{Z}) = \max\{W^R(b,h;\mathcal{Z}) + \sigma_\nu \nu^R, W^D(h;\mathcal{Z}) + \sigma_\nu \nu^D\} \tag{3.14}$$

In (3.14) we add extreme-value taste shocks to the problem. This has the advantage that it smoothes out kinks occurring in the overall value function, thereby greatly facilitating the solution of the problem and improving convergence of the fixed point problems. Yet, there is one subtle computational issue that arises due to taste shocks, namely that households randomize over all feasible choices. This means that default can also be chosen even if the household has positive asset holdings. Since households that are borrowing constrained today are potentially allowed to default next period, in the constrained problem (3.13), the price is not simply the safe price $q_t^S = \frac{1}{R_{t+1}}$. The correct price in this state is a weighted average of the save price and the price faced by a household who chooses to repay, with weights determined by the probability to return to the repayment state. That is, the price in the borrowing constrained problem (3.13) equals $\tilde{q}(b', h) = \lambda q(b', h) + (1 - \lambda)\frac{1}{R'}$.

With extreme-value taste shocks, the expected value of $V(b, h; \mathcal{Z})$ with respect to ν has a closed-form expression given by (McFadden, 1973):

$$\mathbb{E}_\nu \left[V(b, h; \mathcal{Z})\right] = \sigma_\nu \log\left(\exp(W^D(h; \mathcal{Z})/\sigma_\nu) + \exp(W^R(b, h; \mathcal{Z})/\sigma_\nu)\right) \tag{3.15}$$

Similarly, the default probabilities also have a closed-form expression given by:

$$\mathbb{P}(d = 1|b, h; \mathcal{Z}) = \frac{\exp(W^D(h; \mathcal{Z})/\sigma_\nu)}{\exp(W^D(h; \mathcal{Z})/\sigma_\nu) + \exp(W^R(b, h; \mathcal{Z})/\sigma_\nu)} \tag{3.16}$$

3.3.4 Production and price setting

The remainder of the economy is a textbook New Keynesian model as in Bayer et al. (2019). We assume that a mass-zero group of firm managers differentiates an intermediate good Y_t into varieties j, sets prices and distributes all profits to entrepreneur households. The problem these managers face is to maximize the present value of expected real profits given the demand curve for their individual good:

$$\max_{\{p_{jt}\}_{t=0}^\infty} \mathbb{E}_0 \sum_{t=0}^\infty \beta^t \left\{\left(\frac{p_{jt}}{P_t} - MC_t\right) y_{jt} - \frac{\eta}{2\kappa}\left[\log\left(\frac{p_{jt}}{p_{jt-1}}\right)\right]^2 Y_t\right\} \tag{3.17}$$

$$s.t. \quad y_{jt} = \left(\frac{p_{jt}}{P_t}\right)^{-\eta} Y_t \tag{3.18}$$

In equilibrium, all firms behave the same way and charge the same price $p_{jt} = P_t$, which implies from the first-order condition the Phillips curve:

$$\log(\pi_t) = \beta\mathbb{E}_t\left[\log(\pi_{t+1})\frac{Y_{t+1}}{Y_t}\right] + \kappa\left(MC_t - \frac{\eta - 1}{\eta}\right) \tag{3.19}$$

The resulting per-period profits to entrepreneurs are then given by:

$$\Pi_t^f = (1 - MC_t)Y_t - \frac{\eta}{2\kappa}\log(\pi_t)^2 Y_t \tag{3.20}$$

The intermediate good is produced in a competitive environment using labor as the only input and constant returns to scale. They sell this good at real price MC_t

to price setters and solve

$$\max_{N_t} MC_t N_t - w_t N_t \tag{3.21}$$

which implies that wages are equal to real marginal costs:

$$w_t = MC_t \tag{3.22}$$

3.3.5 Government

Lastly, we specify rules for the nominal interest rate and real government debt. Regarding the former, the nominal rate is set according to

$$\frac{R_{t+1}}{\bar{R}} = \left(\frac{\pi_t}{\bar{\pi}}\right)^{\theta_\pi} \tag{3.23}$$

where θ_π captures the degree to which the central bank desires to stabilize inflation at its steady state value $\bar{\pi}$. $\bar{R}$ denotes the steady state nominal interest rate. The fiscal authority behaves according to the following rule:

$$\frac{B_{t+1}}{\bar{B}} = \left(\frac{B_t/\pi_t}{\bar{B}/\bar{\pi}}\right)^{\rho_B} \left(\frac{\pi_t}{\bar{\pi}}\right)^{-\gamma_\pi} \left(\frac{T_t}{\bar{T}}\right)^{-\gamma_T} \tag{3.24}$$

Government expenditures are given by the budget constraint

$$G_t = \frac{B_t}{\pi_t} + T_t - \frac{B_{t+1}}{R_{t+1}} \tag{3.25}$$

where tax revenues $T_t = \tau(w_t N_t + \Pi_t^f + \Pi_t^b)$. The bond rule in (3.24) allows the government to react to business cycle fluctuations. For positive coefficients γ_π and γ_T, which we assume, and increase in inflation or taxes leads to a decrease in government debt, all else equal.

3.3.6 Equilibrium and solution

A recursive competitive equilibrium is a collection of price functions R, w, q, π, policy functions x, b', value functions W^R, W^D, W^C, V, distributions Θ_t, and a law of motion Γ such that:

1. The policy and value functions solve the household problem (3.11) - (3.14).
2. Bond prices solve the intermediary problem (3.9) given the value functions.
3. The nominal interest rate is set according to the Taylor rule (3.23).
4. The government bond rule (3.24) and budget constraint (3.25) are satisfied.

5. The final goods market and labor market clear ((3.19) and (3.22)).

6. The bond market clears:

$$-\frac{1}{R_{t+1}}B_{t+1} = \int q_t(b_{it+1}, h_{it}; S_t)b_{it+1}\mathbb{I}_{f_{it}=0}\mathbb{I}_{d_{it}=0}di + \int \tilde{q}_t(b_{it+1}, h_{it}; S_t)b_{it+1}\mathbb{I}_{f_{it}=1}di \tag{3.26}$$

7. The actual and perceived law of motion coincide, $\Theta' = \Gamma(\Theta, \mathcal{Z}')$.

As is typical of heterogeneous agent models with aggregate fluctuations, the distribution of households is a state variable in the individual households planning problem, and thus the model cannot be solved directly. Instead of applying the algorithm as in Krusell and Smith (1998) to our problem, we employ the linearization technique developed by Bayer and Luetticke (2018) which linearizes the model around its steady state equilibrium without aggregate fluctuations. This amounts to first solving for policy and value functions, bond prices, and the stationary distribution for an environment in which $S_t = 1$ and $\sigma^2_{ht} = \bar{\sigma}^2_h$. The productivity process is discretized into 20 states using the method described in Adda and Cooper (2003), and the asset grid takes 350 unevenly spaced points. We apply the method of endogenous grid points by Carroll (2006) to solve the household planning problem, and augment the algorithm with a maximization step as in Fella (2014) since our problem is not guaranteed to be globally concave. Details on the solution of the household planning problem are provided in Appendix 3.C and Appendix 3.D. Second, to solve for the linearized law motion for states and controls we perform a dimensionality reduction step. For the distribution, we assume fixed copulas for the joint distributions over assets and income conditional on the default flag, with time-varying marginals. The value functions are approximated by sparse polynomials around their steady state values. This has the advantage of considerably reducing the number of state and control variables in the system. With the reduced system, linearization can be performed in the usual way, casting the model in the same form as Klein (2000) and calculating the derivatives with respect to states and controls by finite differences (see also Schmitt-Grohé and Uribe (2004)). Also, it is easy to check whether the local equilibrium is stable and locally unique based on the number of unstable eigenvalues, which is the case in our application.

3.3.7 Calibration

The model is calibrated to the U.S. economy. A period in the model denotes a quarter of a year. Table 3.1 summarizes aggregate steady statistics. The tax rate is

Table 3.1: Steady state statistics

Statistic	Model	Data	Source	Parameter
G/Y (%)	20.01	20	NIPA	Tax rate
B/Y (%)	128.57	180	FRED	W-E transitions
HH Debt/Y (%)	25.64	25.68	FRB	Discount factor
Population filing (%)	0.18	0.21	Livshits et al. (2007)	Default cost
Interest on debt (%)	0.85	3.18	Gordon (2015)	

Notes: We measure B/Y in the data as total public debt less federal debt held by foreign and international investors relative to quarterly GDP between 1997 and 2007. HH Debt/Y is measured as total revolving credit to GDP over the same period. The filing rate, converted from annual to quarterly frequency, comes from Livshits et al. (2007). The interest rate on debt reported by Gordon (2015) is also converted form annual to quarterly frequency. W-E transitions refers to transition probabilities from and to the entrepreneur state.

set to $\tau = 0.208$ so that the ratio of government consumption to GDP equals 20%. We set the discount factor to $\beta = 0.963$ to match total household debt relative to GDP. In the data, household debt is measured as total revolving consumer credit as reported by the Federal Reserve Board (FRB) of Governors, relative to quarterly GDP. Between 1997 and 2007, this ratio was 25.68%. This measure of unsecured credit is also used by Livshits et al. (2007), although their target is slightly higher since they scale credit by disposable income rather than GDP. An issue with this statistic is that consumer credit relative to GDP has grown significantly over time. The time period between 1997 and 2007 features a quiet stable credit-to-GDP ratio and the targeted value is in line with the literature. The model matches this quantity well, although it misses to correctly predict the high cost at which consumers borrow in the data. Further, we target a filing rate in the population of 0.21% every quarter, which corresponds to the average per-household Chapter 7 filing rate as reported by Livshits et al. (2007). Setting the bankruptcy utility cost to $\theta = 5$ delivers a slightly smaller filing rate of 0.18%. In contrast with other work in the literature, we do not explicitly model other sources of consumer bankruptcy such as divorce, medical expense shocks or liability shocks, so our model attributes default entirely to income shocks and randomness in the default decision due to taste shocks. We believe this is not a restrictive assumption, since these additional reasons for bankruptcy are typically modeled as independent from other sources of uncertainty, in particular income risk, and can be thought of as drivers of our taste shocks.

Lastly, using the transition probabilities from and to the entrepreneur state we aim to fit a ratio of government debt to quarterly GDP of 180%. Government debt is calculated as total government debt less federal debt held by foreign and international investors between 1997 and 2007. The calibrated probabilities deliver a ratio equal to 129%. This discrepancy between the model and the data is not too surprising given that the discount rate is low to match the amount of debt held by households in the model. Since we are primarily interested in matching statistics related to consumer bankruptcy, this seems to be an acceptable trade-off.

The remaining parameters are determined as follows. The coefficient of relative risk aversion is set to $\xi = 2$ following Nakajima and Ríos-Rull (2014), and the inverse Frisch elasticity of labor supply to $\gamma = 1$ in line with evidence in Chetty et al. (2011). Following the literature, the quarterly probability regain access to borrowing after default is set to $\lambda = 0.025$, implying that a bad credit record last on average 10 years. The persistence of idiosyncratic worker productivity and the standard deviation of idiosyncratic productivity shocks are taken from Bayer et al. (2019).

Regarding the firm side of the model, we employ standard values from the literature. The elasticity of substitution between varieties $\eta = 20$ implies steady state marginal costs of $MC = \frac{\eta-1}{\eta} = 0.95$ and thus a markup of 5%. Setting $\kappa = 0.0926$ in the Phillips curve implies an average price duration of 4 quarters. In steady state, we assume zero inflation and an annual real interest rate of 2.5%, implying $\bar{\pi} = 1$ and $\bar{R} = 1.0062$ every quarter. The reaction coefficient in the Taylor rule is set to $\theta_\pi = 1.25$ following Kaplan et al. (2018). The parameters governing the dynamics of the aggregate income risk state S_t and reaction coefficients for the government bond rule are taken from Bayer et al. (2019) except for the coefficient ρ_B. We set this parameter to $\rho_B = 0.95$ based on the first-order autocorrelation of the cyclically adjusted time series of real government debt between 1997 and 2007[4]. Table 3.2 provides a complete summary of the model parameters and their values.

3.4 Quantitative results

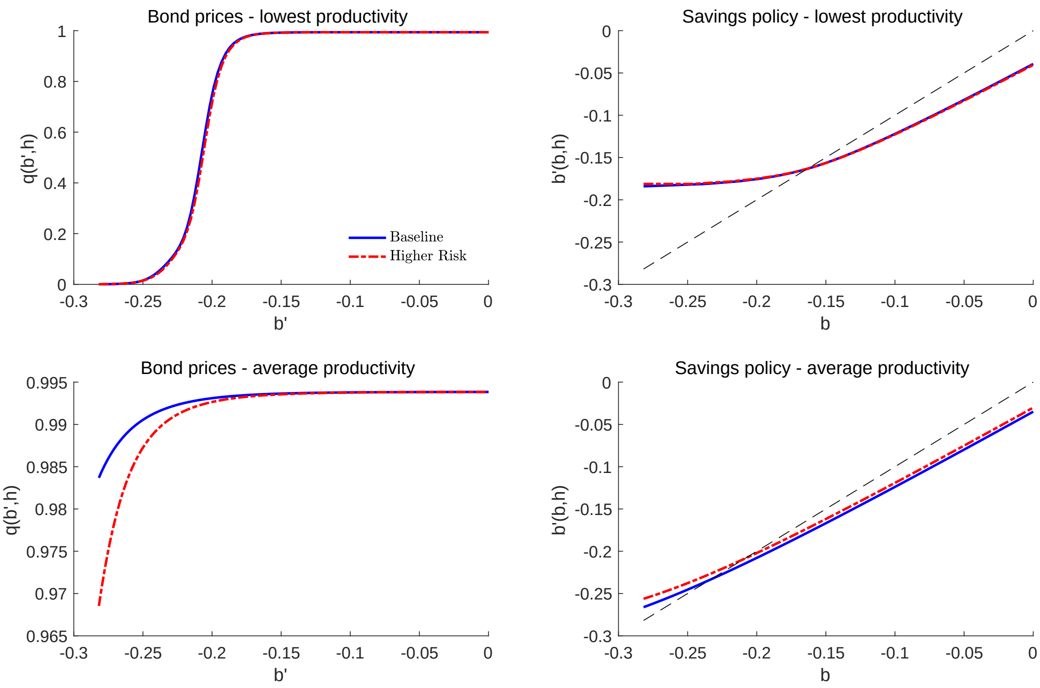
Before we turn to the impact of uncertainty shocks, it is instructive to consider how income risk affects debt prices and default behavior in steady state. For this purpose, we computed another stationary equilibrium of this economy in which income risk is one standard deviation above our baseline calibration, keeping all other parameters the same. Since aggregate variables such as output, wages, and labor supply are completely determined by parameters unrelated to income risk in steady state, this exercise has the advantage that we can clearly isolate how the cost of debt and individual decisions are affected by income risk. Importantly, also the

[4]More precisely, we deflate the nominal government debt series using the GDP deflator, and estimate the cyclical component of the real series in logs using the filter proposed by Christiano and Fitzgerald (2003). Lastly, we estimate an AR(1) process for the cycle between 1997 and 2007 which delivers the coefficient. Over the entire sample, the coefficient is estimated slightly smaller at 0.92.

Table 3.2: Model Parameters

Parameter	Value	Description	Target
Households			
β	0.963	Discount factor	see Table 1
ξ	2	Risk aversion	Nakajima and Ríos-Rull (2014)
γ	1	Frisch elasticity	Chetty et al. (2011)
σ_ν	0.5	Scaling taste shocks	
ϕ	5	Bankruptcy cost	Filing rate 0.21%
λ	0.025	Default flag persistence	10 years
ρ_h	0.98	Autocorr. productivity	Bayer et al. (2019)
$\bar{\sigma}_h$	0.06	STD. productivity shocks	Bayer et al. (2019)
ζ	0.0001	Transition entrepreneur	see Table 1
ι	0.15	Transition worker	see Table 1
Firms			
η	20	Elasticity of substitution	5% markup
κ	0.0926	Price stickiness	4 quarters
Monetary Policy			
$\bar{\pi}$	1	Inflation	0% p.a.
$\bar{R}$	1.0062	Nominal interest rate	2.5% p.a.
θ_π	1.25	Reaction to inflation	Kaplan et al. (2018)
Fiscal Policy			
γ_π	1.5	Reaction to inflation	Bayer et al. (2019)
γ_T	0.5075	Reaction to tax revenue	Bayer et al. (2019)
ρ_B	0.95	Reaction to debt	Autocorrelation of debt
τ	0.208	Tax rate	see Table 1
Aggregate State			
ρ_S	0.84	Autocorr. aggregate state	Bayer et al. (2019)
σ_S	0.54	STD. risk shocks	Bayer et al. (2019)

Figure 3.4: Prices for debt and savings policies across steady states



Notes: Debt prices and savings policies for lowest and average idiosyncratic productivity realizations. Blue lines denote the baseline steady state, red lines the steady state with higher risk.

nominal interest rate is equal in both economies, thus eliminating the endogenous response by the central bank to changes in aggregate demand that will be present when we study impulse responses to income risk shocks.

3.4.1 Steady state comparison

To start, we compare bond prices and savings policies across both steady states. As Figure 3.4 shows, at low productivity realizations, there is little to no change in debt prices consumers face. Consequently, these households do not change their savings behavior as evident from the upper right panel. High debt levels are not chosen by these households as the default premium on such loans drives the debt price toward zero in equilibrium, leading to the savings policy becoming flat in the high debt region.

For comparison, we also show bond prices and savings policies for average productivity households in the lower two panels. In the higher risk steady state, loans with a higher debt balance are more expensive since it is more likely to experience income realizations that lead to default by the borrower. The lower right panel shows that higher income risk leads to an increase in savings, which is to be expected given the precautionary motives at play in this type of economy. Yet, we clearly observe that part of this is driven by deleveraging of indebted households, and particularly of highly indebted ones. This is because not only do households want to save more for precautionary motives, but it is also the endogenously determined borrowing limits that become tighter.

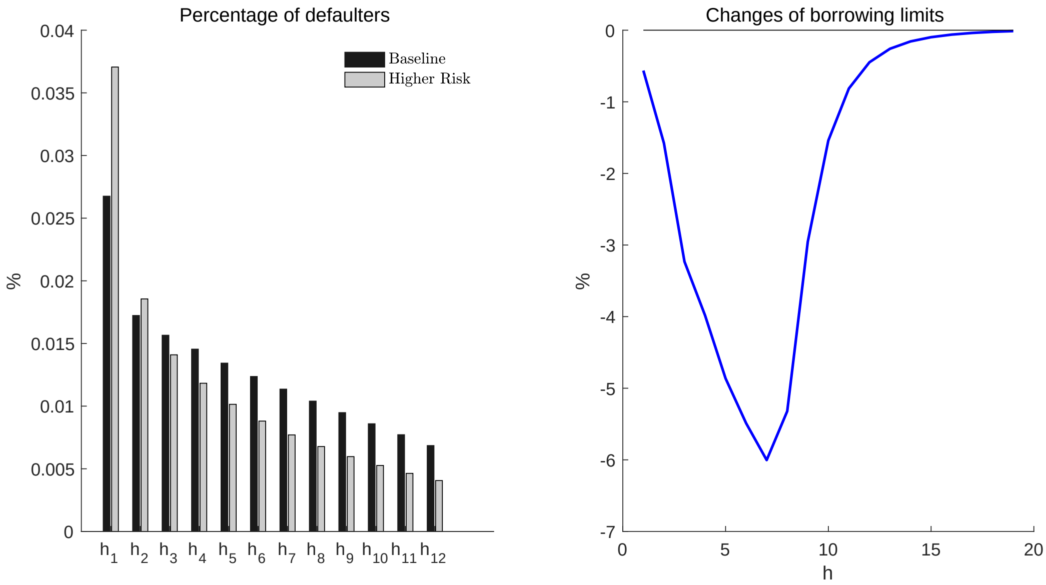
The borrowing limit here denotes the largest debt balance a borrower can optimally choose given the bond prices he faces, and it is defined as (recall that we define $b_{i,t} < 0$ as debt):

$$\bar{b}_{it+1} \equiv \min_{b_{it+1}} q(b_{it+1}, h_{it}; S_t) b_{it+1} \tag{3.27}$$

It is well known that the amount promised, $b_{i,t+1}$, and the amount of resources a borrower can obtain, $q(b_{it+1}, h_{it}; S_t)b_{it+1}$, follow a "Laffer Curve" relationship in a setup such as ours (e.g. Arellano (2008)). For small debt levels, the default probability is zero, and the amount of resources received for a given promise depends only on the safe interest rate. As the debt level increases, so does the default probability, and the bank becomes increasingly reluctant to lend to the household. At some point, the default probability is high enough to cause the amount of resources the bank is willing to transfer to the household to actually decline by virtue of (3.9). The point where this happens is precisely $\bar{b}_{it+1}$. This implies that debt levels larger than $\bar{b}_{it+1}$ will optimally not be chosen, since for such debt levels, there exist smaller loans through which the consumer receives the same amount of resources today and incurs a smaller liability tomorrow. Clearly, since loans are priced competitively, there is such a debt limit for each productivity realization. As it turns out, the effect of idiosyncratic income risk on this individual debt limit is highly non-linear and non-monotone in individual productivity.

This is shown in the right panel of Figure 3.5 where we show the percentage change in the borrowing limit across steady states conditional on the current productivity state. Borrowing constraints tighten most strongly for the bottom half of the productivity realizations, reaching a roughly 6% decline in the borrowing limit. This affects mostly households with below average but not the lowest productivity, while very lowly and very highly productive households see little change in their ability to borrow. This is intuitive, as low income households have high default risk and thus a small menu of debt contracts available to begin with, and high income households have little incentive to default either way, leaving bond prices and thus borrowing limits at the extremes largely unaffected. Around the middle of the income distribution however, higher income risk, and thus higher default risk, decreases bond prices significantly. Lastly, the left panel of Figure 3.5 shows that default rates in the model are predominantly driven by very low income households, and that these households in fact default more often in the higher risk steady state,

Figure 3.5: Default rates and borrowing limits across steady states



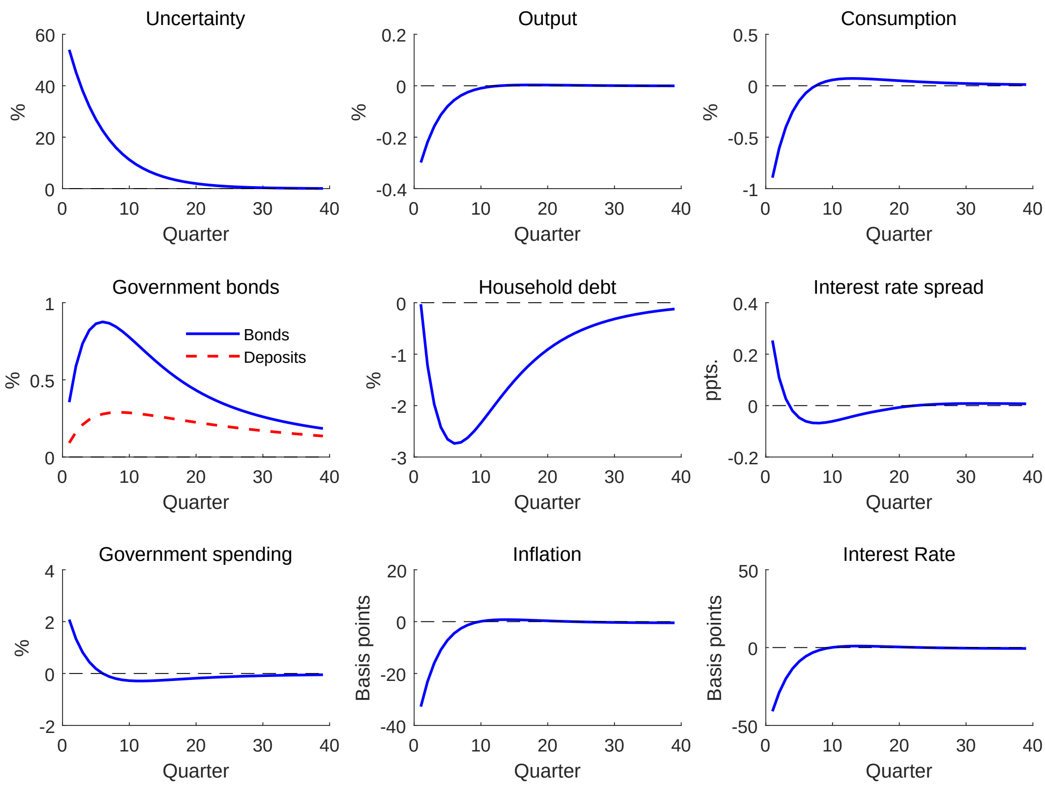
Notes: Left panel: Percentage of households that default with a particular productivity realization in both steady states. Right panel: Percent change of endogenous borrowing limit relative to baseline steady state.

while default rates in the rest of the income distribution fall, consistent with higher savings by these households.

3.4.2 Effects of income risk shocks

We now turn to the aggregate effects of income risk shocks. In Figure 3.6 we show aggregate responses to one standard deviation shock to household income risk. In line with the empirical results, aggregate output and consumption decline in response to an increase in aggregate income risk. The output response is similar in magnitude to that observed in the data, while consumption declines by 0.9% on impact. The increase desire in savings by households, which correspond to deposits in the banking sector, causes an increase in government debt and decrease in household debt (second row). The drop in aggregate demand causes prices to fall drastically, in response to which the central bank cuts interest rates. The government also increases its consumption to stabilize the economy, thus offsetting some of the decline in private demand. The response of household debt matches well the empirical response both in terms of duration and magnitude. At the trough, household debt is around 2.8% below its steady state value, as default rates increase and households increase savings to safeguard against future income shocks. The interest rate spread, defined as the difference between the average interest rate on debt minus the nominal interest rate, goes up by 25 bps. on impact. This eats up more than 50% of the monetary

Figure 3.6: Aggregate responses to income risk shock

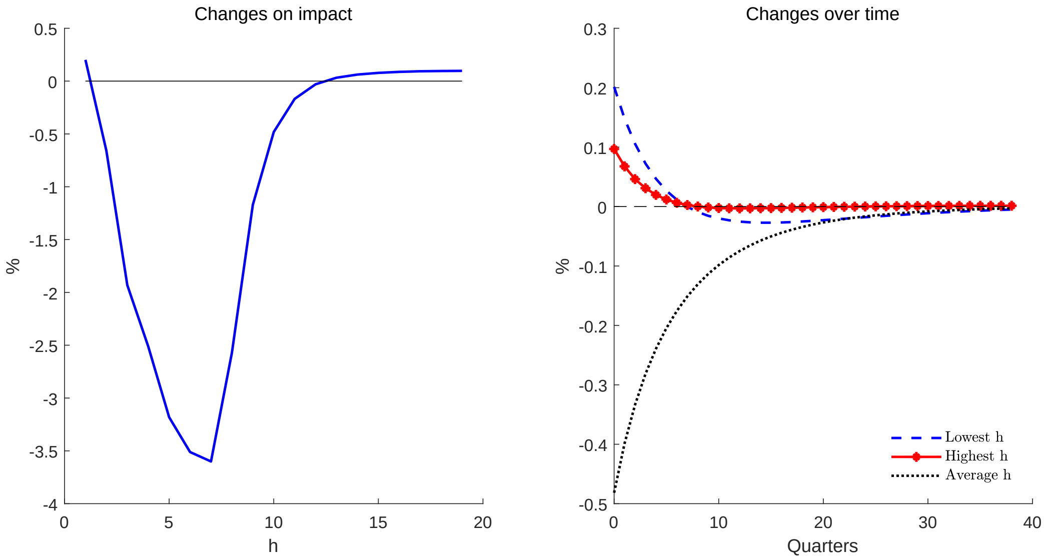


Notes: Impulse responses to a one standard deviation increase in aggregate income risk. The interest rate spread is calculate as the difference between the average interest rate on debt and the nominal interest rate. The responses of the spread, inflation, and the interest rate are annualized.

stabilization for indebted households in terms of the interest rate they face. Yet, the model somewhat misses the magnitude in the spread response (0.25 ppts. in the model vs. 0.5 ppts. in the data) and the slow buildup in the spread that we observed in the empirical impulse response.

Turning next to the effects of income risk shocks on borrowing constraints, in Figure 3.7 we show in the left panel the immediate impact of an uncertainty shock on borrowing limits across individual productivity states. We observe the same non-monotonic relationship we found in the steady state comparison, although the effect is less pronounced. This result comes from two sources. First, in contrast to the steady state analysis, where income risk was permanently higher, income risk shocks now die out eventually. Second, in response to the decline in aggregate demand, the central bank cuts interest rates, which ceteris paribus increases bond prices and partly offsets the risk-induced decline in borrowing limits. Also noteworthy is the small expansion of the borrowing limit for very low-income households. This is due to the fact that households with low initial productivity actually expect their productivity to increase in the future. The converse is true at the upper end of

Figure 3.7: Changes in borrowing limits



Notes: Left panel: Percentage change in borrowing limit in the shock period, conditional on idiosyncratic productivity. Right panel: Change in borrowing limit over time for productivity realizations.

the productivity distribution, where individual productivity is expected to fall in the future. However, as we discussed earlier, these households rarely ever default anyway, so the change in the borrowing limit is driven entirely by the interest rate response. As shown in the right panel of Figure 3.7, the change in borrowing limits persists over time. For households with the lowest or highest productivity level, borrowing limits increase relative to steady state for about a year, and this increase is considerably larger for very lowly productive households. In contrast, borrowing limits of average productivity households are persistently depressed after an income risk shock, and revert back to steady state very slowly over time.

3.4.3 Decomposing the transmission through consumer default

The results from the previous section are driven by several important channels. First, banks anticipate higher default rates due to higher income risk. Thus, banks raise interest rates to compensate for this higher risk, which in turn lowers consumption. Second, households also increase savings for precautionary motives since they expected larger dispersion in future incomes and have prudence as part of their utility. In addition, a third and more subtle channel is that the curvature in the interest rate schedule by itself leads to an increase in expected interest rates due to higher risk, even under the stationary equilibrium schedule.

To better understand how these different channels contribute to the transmission of shocks, it is helpful to consider the consumption Euler equation that needs to hold when households choose to repay:

$$u'(x_{it}) = \frac{\beta}{q(b_{it+1}, h_{it}; S_t) + b_{it+1}\frac{\partial}{\partial b_{it+1}}q(b_{it+1}, h_{it}; S_t)}\mathbb{E}_{\mathcal{Z}}\left[\frac{u'(x_{it+1})}{\pi_{t+1}}\mathbb{P}(d_{it+1} = 0|b_{it+1}, h_{it+1})\right] \tag{3.28}$$

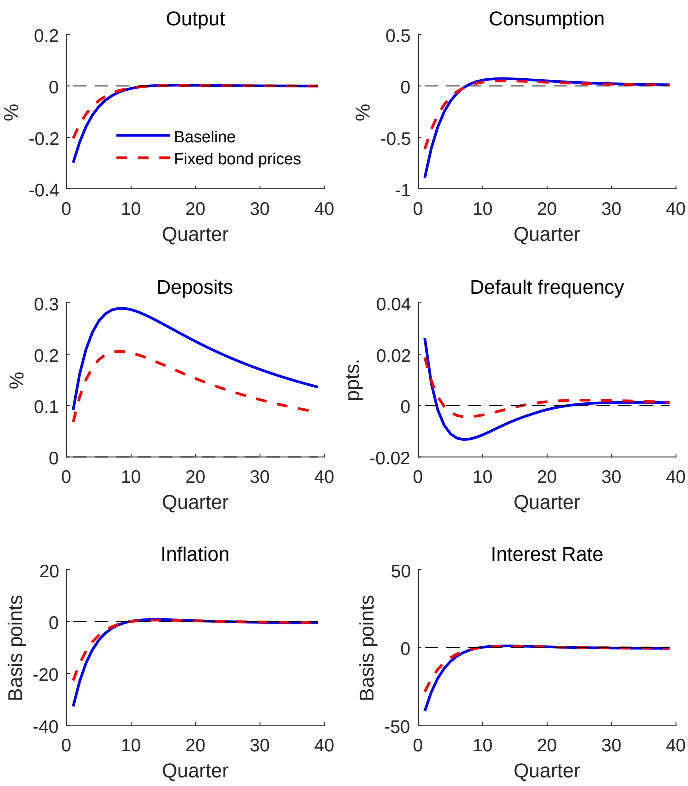
Precautionary savings are captured by the expected real marginal utility in the next period. Similar to standard consumption Euler equations, the first fraction on the right-hand side can be interpreted as the discounted nominal interest rate for a loan of size b_{it+1}. It is composed of the marginal increase in resources given the default risk plus the increase (decrease) in the risk premium that is anticipated due to an increase (decrease) in loan size. This term captures the anticipation effect by banks. In a world in which consumers never default, this anticipation effect is absent and the term in the denominator collapses to the inverse of the nominal interest rate set by the central bank.

In order to assess the importance of the anticipation mechanism in driving the baseline results we perform the following experiment. Suppose that in response to an income risk shock, banks do not change their prices and behave as if nothing had changed relative to steady state. This means that banks do not anticipate that the exogenous transition probabilities between productivity states have change, and they also do not anticipate changes in behavior in future periods[5]. The comparison in Figure 3.8 shows that relative to the baseline results, the impact response of output and consumption is muted when bond prices are fixed at their steady state values. We now find a roughly 0.2% decline in output, and a 0.6% decline in consumption, which implies that roughly a third of the baseline responses of output and consumption are driven by the fact that banks tighten lending conditions. Since prices do not change when households make their decisions, in the aggregate they accumulate less savings following an income risk shock. This is visible both in the smaller increase in government debt and the smaller increase in deposits in the banking sector. As expected, aggregate default rates are smaller on impact relative to the baseline results, since the ability to borrow is not limited by the endogenous change in bond prices in this experiment. This causes a smaller drop in bank profits on impact, but the incorrect pricing of consumer credit in future periods leads to persistent losses by banks.

These results have two interesting implications. First, the fact that banks tighten credit constraints for consumers is a quantitatively important driver of the economic decline following an increase in idiosyncratic income risk. Second, the fact that the output and consumption drops are still sizable even when this effect is shut down

[5] This experiment requires to redefine profits in the banking sector to also include all resources transferred to households and the government between t and $t + 1$.

Figure 3.8: Aggregate responses with fixed bond prices



Notes: Solid lines denote the baseline responses, dashed lines the responses when bond prices used by banks remain at their steady state values.

suggests that the option to default does not play a stabilizing role for business cycles, but rather the opposite. This stands in contrast to recent work by Auclert et al. (2019) where ex-post debt forgiveness stabilizes economic activity. Our results mirror those in Nakajima and Ríos-Rull (2014) where higher default premia during recessions deter household borrowing, thus making aggregate consumption more volatile, not less, relative to the no default case.

3.5 Monetary policy

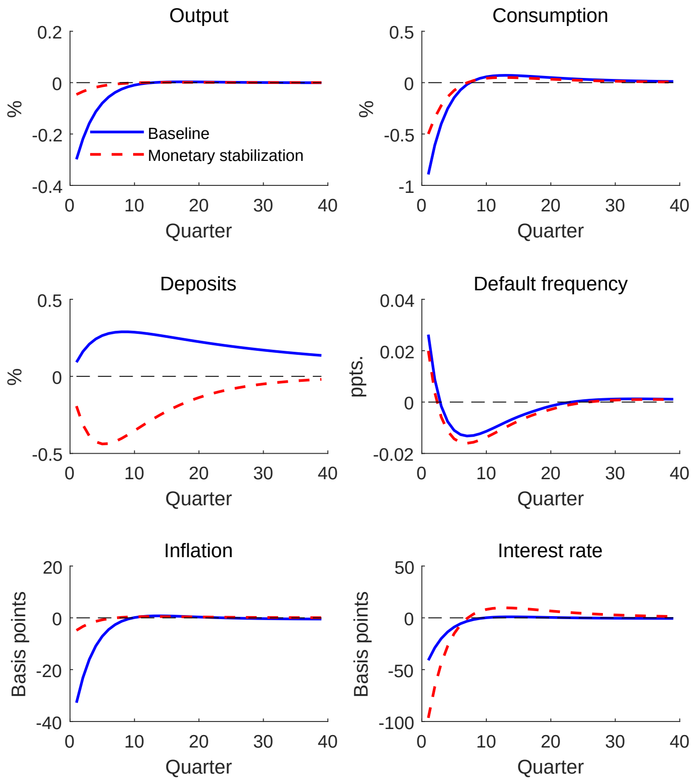
We now consider in more detail the effects of monetary policy. Specifically, we first ask how monetary policy shapes aggregate outcomes when it systematically reacts more aggressively to inflation. Secondly, we study the role of monetary policy shocks in our environment.

3.5.1 Systematic monetary policy

The central bank directly affects borrowing constraints in our economy since debt is priced according to the risk adjustment on top of the savings price. As such, by cutting interest rates, bond prices increase ceteris paribus, which works against the downward adjustment in bond prices due to higher risk. To get a better understanding of this opposing effect, we study a situation in which the central bank reacts much more strongly to deviations of inflation from target by setting $\theta_\pi = 20$, which we call "monetary stabilization". All other parameters remain unchanged. Figure 3.9 shows the results for key aggregate outcomes. The impact drop in output, consumption, and inflation are less pronounced, because the central bank strongly discourages household saving. Relative to steady state, consumption falls by 0.5% instead of 0.9% in our baseline model, and the output decline is minuscule compared to our baseline results. The nominal interest rate declines by almost 100 basis points annually on impact, more than twice the baseline response, and the significantly smaller decline in aggregate demand naturally translates into a much smaller drop in prices. The strong decline in interest rates leads households to actually dissipate savings, shown by the negative response of deposits, and the muted increase in government debt. We also observe that part of the aggregate stabilizing effect works through lower default rates. On impact, default rates increase by less compared to the baseline economy, and default occurs less frequently over time as well. The smaller impact response of default rates then translate into reduced losses by banks on impact. Thus, by implementing a very aggressive policy, the central bank can significantly dampen the economic decline on impact.

As we discussed earlier, the central bank has some control over individual bond prices since it can change the discount rate that applies uniformly to all loans. The effects on borrowing limits are shown in the left panel of Figure 3.10. Since the central bank cannot affect individual bond prices but only the relevant discount rate, the borrowing limit moves up rather uniformly. For low income households, monetary policy is able to engineer a larger expansion in the credit limit on impact, which comes from the previously noted expected growth in productivity for these households and the now larger drop in the interest rate. This relative credit expansion leads to greatly reduced overall default rates in the economy on impact and a protracted decline in default rates over time. Hence, monetary policy can lean against the economic decline and tightening of borrowing constraints by forcefully reducing the policy rate. At the high end of the income distribution, the massive interest rate cut deters savings, which also contributes to a smaller economic decline as these households dissave through increased consumption. However, Figure 3.9 reveals that the nominal rate overshoots the steady state in the medium term. This has the effect that over time, borrowing limits contract again. Since borrowing limits

Figure 3.9: Aggregate effects of monetary stabilization

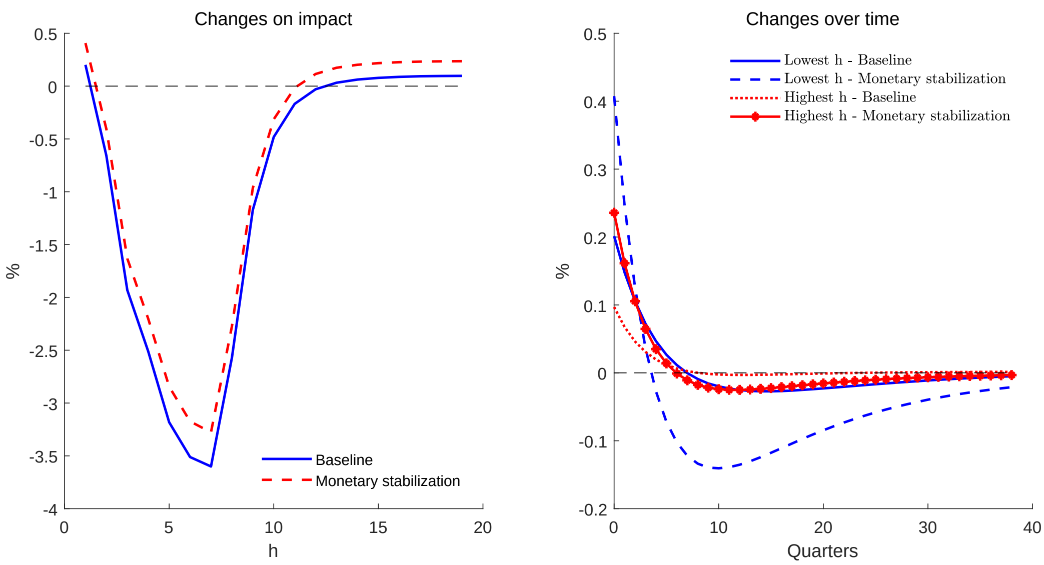


Notes: Impulse responses to a one standard deviation increase in aggregate income risk. Solid lines show the baseline response, dashed lines the response when $\theta_\pi = 20$. The responses of inflation and the interest rate are annualized.

are more sensitive to changes in interest rates for low-income households, they see a large contraction in their ability to borrow over time, as the right panel of Figure 3.10 illustrates.

The fact that the nominal interest rate overshoots the steady state in the medium run is due to the fact that under monetary stabilization the household sector in the aggregate accumulates more debt than in the baseline model, as can be seen from the government debt response as well as the default response. But what this implies over the medium to long term is that the central bank also has to bring back household savings and debt back to steady state levels. The central bank achieves this by raising interest rates at a much faster pace. Thus, the central bank initially creates more room for household indebtedness, which later on has to be reversed the same way it was created, by a forceful adjustment in the interest rate. As shown in Figure 3.10, this brings about large fluctuations in borrowing limits especially for rather poor households. At the same time, poor households are likely to benefit the most from the initially smaller decline in the borrowing limit. Hence, it interesting to consider the consumption response of households in the cross section conditional

Figure 3.10: Borrowing limits under monetary stabilization

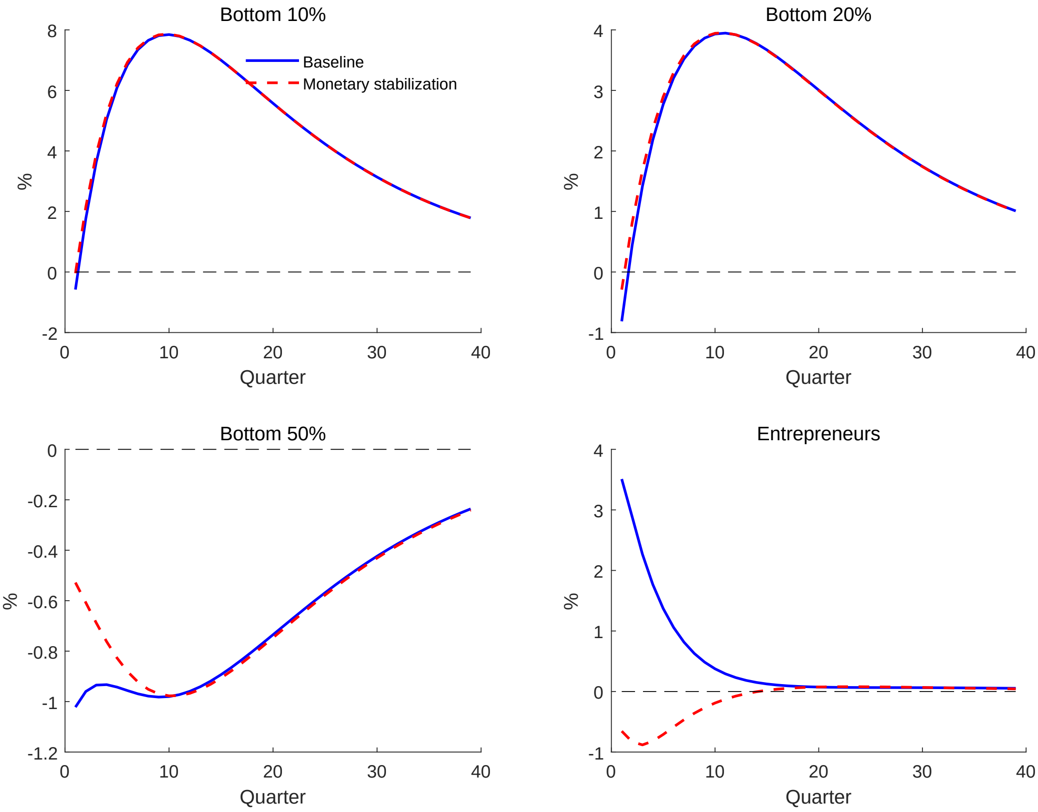


Notes: Left panel: Percentage change in borrowing limit in the shock period, conditional on idiosyncratic productivity. Solid lines show the baseline response, dashed lines the response when $\theta_\pi = 20$. Right panel: Change in borrowing limits over time for different labor productivity realizations.

on the position in the income distribution. We calculate the impulse responses of consumption for households that hold debt across the different productivity levels, which, given the assumed preferences of consumption and leisure, is synonymous with different income levels.

The results are shown in Figure 3.11. Starting in the upper left panel, we see a slight expansion in consumption by very low income households when monetary policy strongly stabilizes, but overall the differences between the responses are minor, and the fact that the borrowing limit for low income households fluctuates does not lead to their consumption becoming more volatile. The upper right panel widens the group of households to cover the bottom 20% of the income distribution. Clearly, the pattern is very similar to the bottom 10%, although smaller in magnitude overall. This suggests that the difference in the aggregate consumption response is driven by households toward the middle of the income distribution. The bottom left panel shows this, where we further widen the group of households to include the bottom half of the income distribution. First, contrary to the response of very low income households, those households closer to median income actually see an overall decline in consumption relative to steady state. Second, under monetary stabilization, the initial decline in consumption is about half of the response under the baseline monetary regime. This means that around the middle of the income distribution, monetary policy has the most bite in relaxing borrowing constraints, but

Figure 3.11: Consumption responses by position in income distribution



Notes: Percentage changes of consumption relative to steady state conditional on households belonging to the respective quantile of the income distribution in a given quarter after the shock. The last panel shows the response of entrepreneurs for comparison. Solid lines show the baseline responses, dashed lines the responses when $\theta_\pi = 20$.

this obviously comes at the cost of redistributing from high income to low income households. As the bottom right panel shows, while entrepreneurs, the savers in the economy, increase consumption in the baseline economy, they actually decrease consumption when the central bank strongly seeks to stabilize inflation. The reason is that these households want to increase their savings in response to higher income risk, but the forceful reduction in the interest rate increase the price for savings by so much that the desired buffer stock savings can only be financed by cutting back on consumption.

3.5.2 Monetary policy shocks

Next to the systematic component of monetary policy, a second natural question to ask is what the effects are of exogenous monetary policy shocks in our environment. We augment the Taylor rule in (3.23) by an autocorrelated shock, such that the

Taylor rule now becomes

$$\frac{R_{t+1}}{\bar{R}} = \left(\frac{\pi_t}{\bar{\pi}}\right)^{\theta_\pi} e^{\epsilon_t^M} \tag{3.29}$$

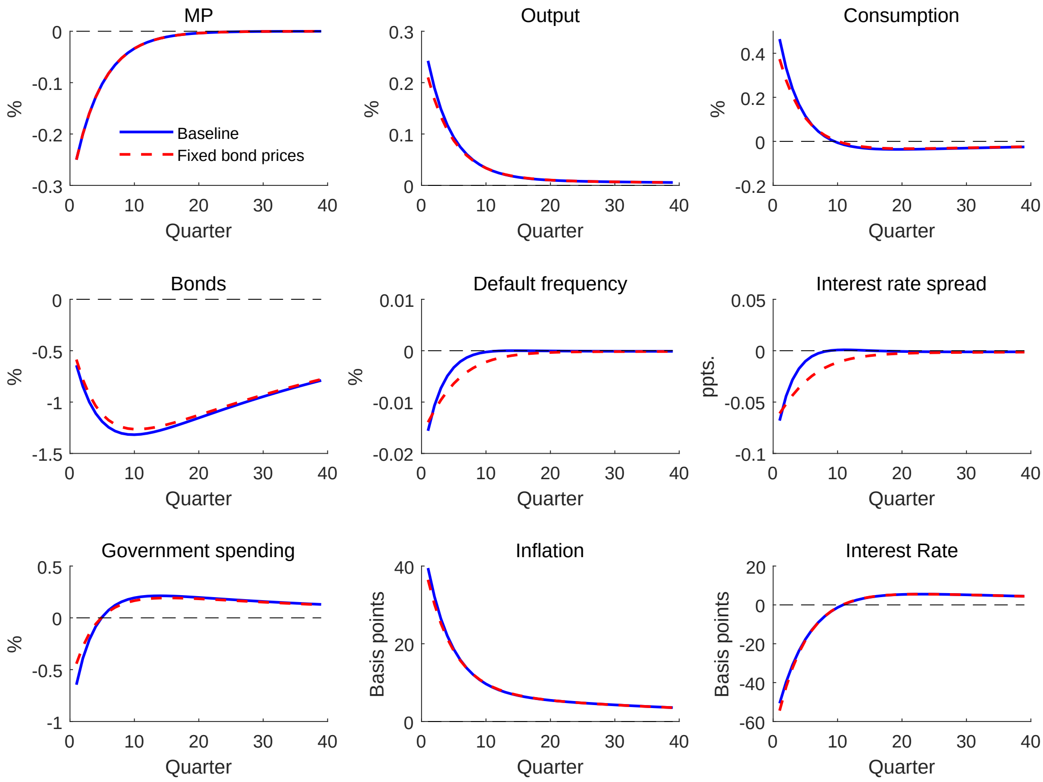
$$\epsilon_t^M = \rho_M \epsilon_{t-1}^M + v_t \tag{3.30}$$

For the quantitative analysis we set $\rho_M = 0.8$ and consider an expansionary shock that, all else equal, lowers the nominal interest rate by 25 basis points on impact.

Monetary policy shocks have, of course, the immediate effect of expanding borrowing limits of every household and making saving less attractive. This leads to an increase in consumption, and hence output, but also to a reduction in overall default rates as rising income and easier access to credit increase a household's ability to repay debt. As the shock is known to have some persistence, incomes are expected to be higher in the future as well. This in turn leads to lower default probabilities next period, and this reduction feeds back into bond prices at the time the shock occurs, expanding borrowing limits even further. This novel feedback effect is absent in models without the default option, and it is interesting to investigate how it reinforces the effects of monetary policy shocks. To do so, as before, we fix default probabilities at their steady state values, thus eliminating the second-round feedback effects.

Figure 3.12 shows the results. Consistent with the reasoning above, the output and consumption response are smaller compared to the baseline economy when the second-round feedback effects are turned off. As a results, the initial increase in inflation and the resulting offsetting effect on the nominal interest rate are larger. Comparing the output and consumption responses across the two scenarios, we find a roughly 15% larger output response and a 25% larger consumption response when banks correctly price consumer credit. Furthermore, households decrease savings to a lesser extend, and default rates and the interest rate spread turn out smaller. The changes in default rates and interest rate spreads take longer to revert back to steady state when default probabilities are fixed because a smaller initial increase in household debt, resulting from the missing second-round expansion in borrowing limits, implies lower default incentives and thus smaller default rates. Thus, also for the case of monetary policy shocks we find that the anticipation of future default rates by banks amplifies the response of the economy as a whole.

Figure 3.12: Aggregate responses to expansionary monetary policy shock



Notes: Impulse responses to an expansionary monetary policy shock which, all else equal, reduces the nominal interest rate by 25 basis points on impact. Solid lines show responses in the baseline economy, dashed lines the response when bond prices are fixed at their steady state values. The responses of inflation and the interest rate are annualized.

3.6 Conclusion

In this paper we investigate the macroeconomic effects of changes in household income risk in an economy with consumer bankruptcy and nominal rigidities. Consistent with empirical evidence, the model predicts that higher uncertainty tightens borrowing constraints that households face, which leads to amplified business cycle fluctuations relative to a model where banks to not anticipate future default risk. As households not only increase savings after an increase in income risk but also strongly reduce their debt levels, aggregate demand-driven drops in output and consumption become larger. Our results emphasize that the possibility to default introduces higher sensitivity of the cost of credit to changes in income risk, which amplifies business cycle movements. As such, in a demand driven economy, the economy is less resilient to risk shocks despite the state-contingency inherent in defaultable debt. We also show that the default anticipation channel aggravates the effects of other shocks, such as monetary policy shocks, by roughly 20%.

Monetary policy can counteract the economic decline by drastically cutting interest rates, thereby relaxing borrowing constraints, but this inevitably leads to monetary policy have to reign in household debt through forceful hikes in interest rates later on, which creates large swings in borrowing constraints. The transmission of monetary policy shocks works both through a direct expansion of credit limits and a feedback effect from lower default rates on credit prices, but the latter effect seems quantitatively unimportant, at least in the aggregate.

There are several aspects which are worth addressing in future work. First, we have focused on a setup that features only a single asset through which consumers can insure themselves against risk. Extending the model to allow for durable goods and both unsecured and secured credit as in Hintermaier and Koeniger (2016) seems natural as this offers a more detailed description of household portfolios. Also, in models with both liquid and illiquid assets, an increase in income risk leads to portfolio rebalancing towards more liquid assets as in Bayer et al. (2019) or Guerrieri and Lorenzoni (2017), which may interact in interesting ways with the change in access to unsecured credit that we study.

Second, we model consumer bankruptcy in a very stylized way in that we abstract from additional issues regarding personal bankruptcy such as delinquencies or debt restructuring. Allowing for a more detailed description of bankruptcy regulation can shed more light on how different bankruptcy schemes impact the channel in this paper, especially since there have been significant changes in regulation over time. For example, a "No Fresh Start" system along the lines of Livshits et al. (2007) would provide an extension to our model that is closer to a different bankruptcy scheme actually in practice. Since the structuring of the bankruptcy code impacts the way unsecured debt is priced, it would be interesting to study how different systems mitigate or reinforce the effect of income risk on credit prices.

APPENDICES

3.A Data description

This section provides more information on the data used for estimating impulse responses and calibration of the model. The data used in Section 3.2 and Appendix 3.B mostly come from the FRED database, with the exception of time series of the level of income risk and income risk shocks. The following table summarizes sources and units of the original data. Most of these series warrant no further explanation. The charge-off rate measures the value of credit card loans removed from the books of all commercial banks, net of any recoveries and relative to average loan size. Credit card interest rates measure the average interest rate charged on credit card plans. This rate and the federal funds rate are averages of monthly values in a respective quarter. Consumer credit denotes the nominal level of outstanding liabilities of households and nonprofit organizations, which we deflate using the GDP deflator series.

Table 3.A.1: Data overview

Variable	Source	Code
Real GDP	FRED Database	GDPC1
GDP Deflator	FRED Database	GDPDEF
Real Consumption	FRED Database	PCECC96
Real Investment	FRED Database	GPDIC1
Unemployment rate	FRED Database	LRUN64TTUSQ156S
Effective Federal Funds Rate	FRED Database	FEDFUNDS
Charge-Off Rate	FRED Database	CORCCACBS
Credit Card Interest Rate	FRED Database	TERMCBCCALLNS
Lending Standards	FRED Database	DRTSCLCC
Consumer Credit	FRED Database	CCLBSHNO
Income Risk Level	Bayer et al. (2019)	
Income Risk Shocks	Bayer et al. (2019)	

Lending standards refers to the net percentage of domestic banks having tightened their lending standards or terms on credit card loans. This data is based on the "Senior Loan Officer Opinion Survey on Bank Lending Practices". The July 2018 release of this survey (Federal Reserve Bank, 2018) describes questions regarding lending standards as follows:

For questions that ask about lending standards or terms, ”net fraction” (or ”net percentage”) refers to the fraction of banks that reported having tightened (”tightened considerably” or ”tightened somewhat”) minus the fraction of banks that reported having eased (”eased considerably” or ”eased somewhat”).

3.B Additional empirical results

Here we provide robustness checks for the Local Projections from section 3.2. As before, y_t denotes 100 times the natural log of the level of a given variable, except for interest rates, the unemployment rate, and the net percentage of banks increasing lending standards. As a first robustness check, for real GDP, consumption, and consumer credit we redefined the dependent variable as the cumulated growth rate between $t-1$ and $t+h$ and estimated impulse response functions for these variables from the following model:

$$\Delta_h y_t = \alpha_h + \theta_h \epsilon_t + u_{t+h} \tag{3.31}$$

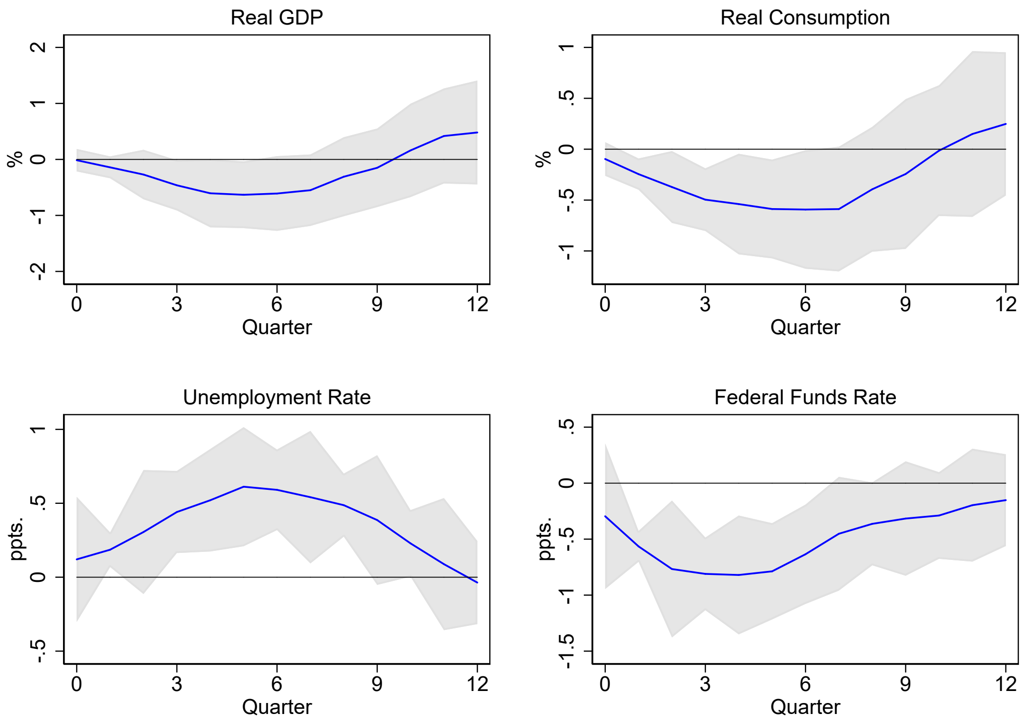
where $\Delta_h y_t \equiv y_{t+h} - y_{t-1}$. For all other variables, the model is the same as in section 3.2 except that we drop past values of the dependent variable. Figures 3.B.1 and 3.B.2 below show that the results are virtually identical to our baseline results.

Next, we include a rich control set into the Local Projections. The model then reads

$$y_{t+h} = \alpha_h + \theta_h \epsilon_t + \Gamma_h X_{t-1} + u_{t+h} \tag{3.32}$$

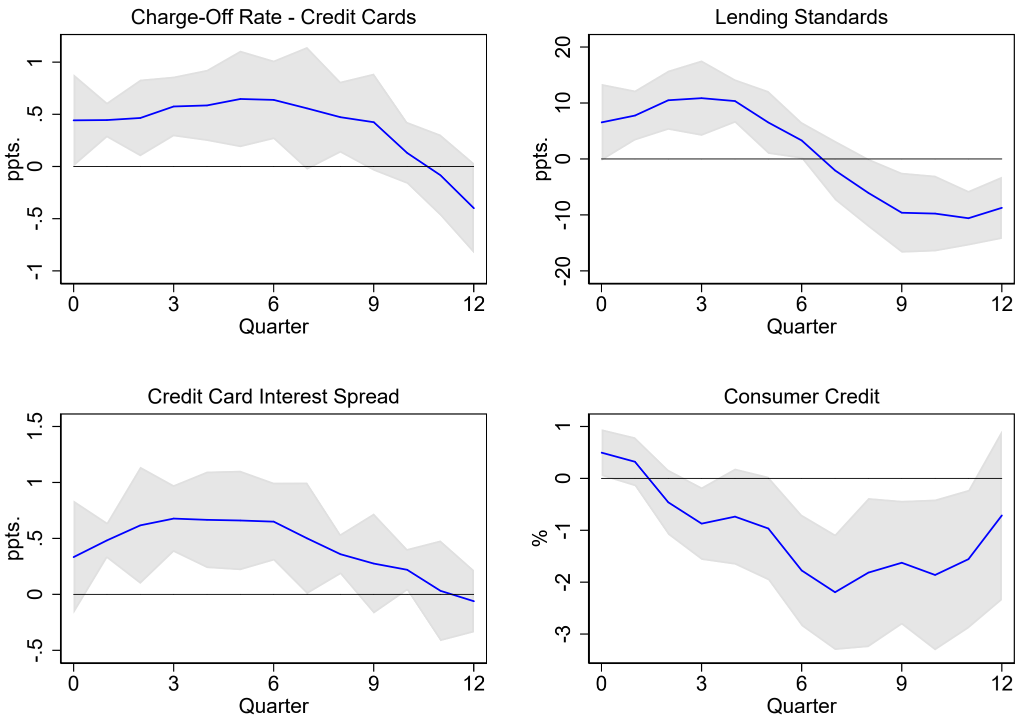
Specifically, we include in X_{t-1} one lag of GDP, consumption, investment, consumer credit (all in real terms), the unemployment rate, the effective federal funds rate, the GDP deflator, chargeoff rates, the interest rate spread on credit card loans, and the percentage of banks increasing lending standards. Finally, we include one lag of the estimated time series for the risklevel and the risk shocks from Bayer et al. (2019). Figures 3.B.3 and 3.B.4 show that the results are also robust to this change, if not stronger.

Figure 3.B.1: Aggregate responses to income risk shocks - cumulated first differences



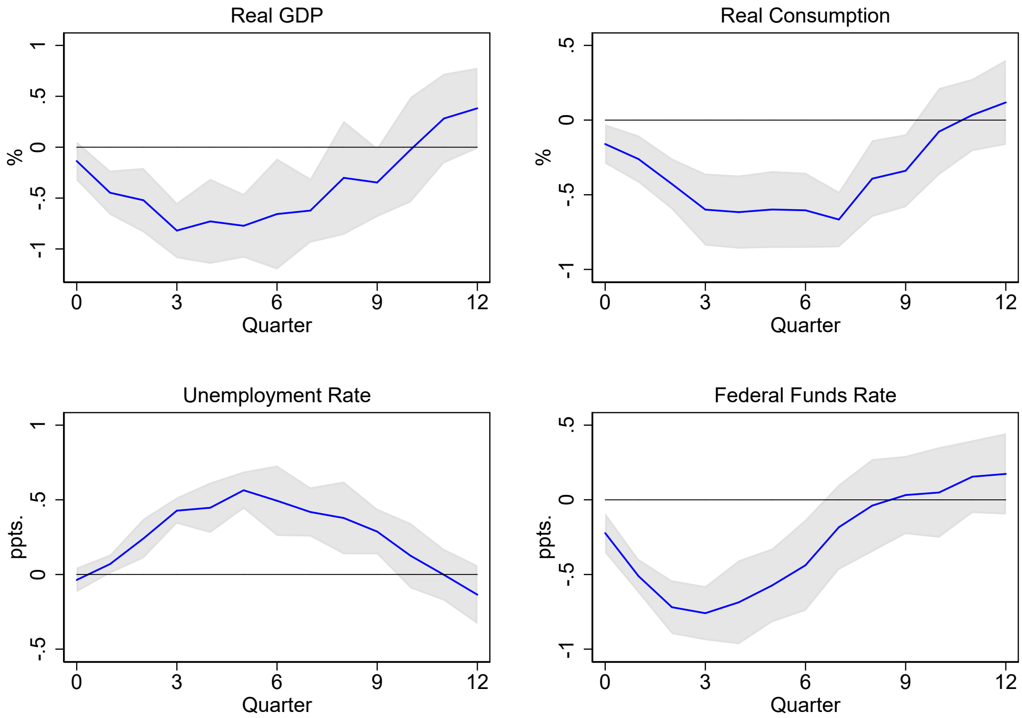
Notes: Income risk shocks come from Bayer et al. (2019), the remaining series come the FRED database. Shaded areas denote 90% error bands constructed using Newey-West standard errors. Impulse response functions estimated without controls and cumulated first differences of the dependent variable.

Figure 3.B.2: Responses of financial variables to income risk shocks - cumulated first differences



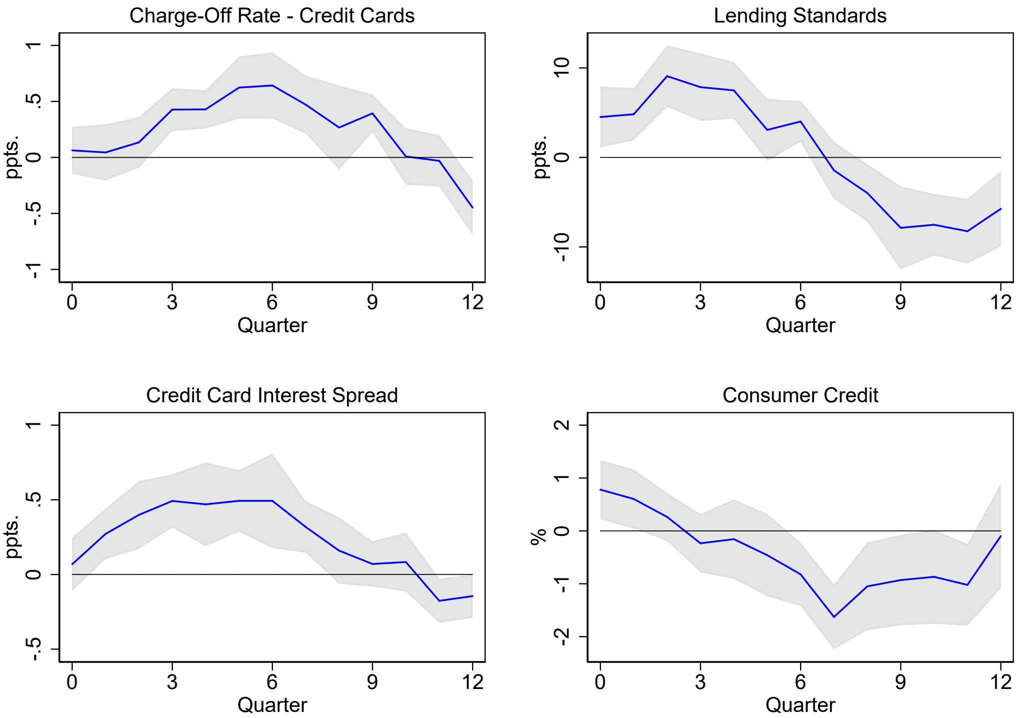
Notes: Income risk shocks come from Bayer et al. (2019), the remaining series come the FRED database. Shaded areas denote 90% error bands constructed using Newey-West standard errors. Impulse response functions estimated without controls and cumulated first differences of the dependent variable.

Figure 3.B.3: Aggregate responses to income risk shocks - with controls



Notes: Income risk shocks come from Bayer et al. (2019), the remaining series come the FRED database. Shaded areas denote 90% error bands constructed using Newey-West standard errors. Impulse response functions estimated in levels with controls.

Figure 3.B.4: Responses of financial variables to income risk shocks - with controls



Notes: Income risk shocks come from Bayer et al. (2019), the remaining series come the FRED database. Shaded areas denote 90% error bands constructed using Newey-West standard errors. Impulse response functions estimated in levels with controls.

3.C Monotonicity of savings policies

Here we provide the formal argument that the optimal savings policy is a monotone function of beginning-of-period assets, conditional on a given productivity level. Clearly, the non-concavity only occurs when a household has the option to default. To establish monotonicity we apply Proposition 4 in Gordon and Qiu (2018) to our problem. The argument is identical to their proof of policy function monotonicity in the Arellano (2008) model. To save on notation, we omit the aggregate states as arguments from the value functions, since these only play a role outside the steady state without aggregate fluctuations. First, we define the optimal savings policy as

$$G(b,h) \equiv \underset{b' st.x>0}{\arg\max} \left\{ u(x(b',b,y(h))) + \mathbf{W}(b',h) \right\} \tag{3.33}$$

where

$$x(b',b,y(h)) = \frac{b}{\pi} + y(h) - q(b',h)b' \tag{3.34}$$

$$y(h) = (1-\tau)\left(\frac{\gamma}{1+\gamma} whN(w) + \mathbb{I}_{h=0}\Pi\right) \tag{3.35}$$

$$\mathbf{W}(b',h) = \beta\mathbb{E}_{h'|h}\left[\sigma_\nu \log\left(e^{W^D(h')/\sigma_\nu} + e^{W^R(b',h')/\sigma_\nu}\right)\right] \tag{3.36}$$

This definition simply rewrites the problem in terms of the current period payoff and the continuation value in more general form. A prime denotes next period variables. Then, we need to check that the following three conditions are satisfied:

1) $x(b',b,y(h))$ is increasing in b.
2) $x(b',b,y(h))$ has increasing differences in b, b'.
3) $\mathbf{W}(b',h)$ is increasing in b'.

The first condition is clearly satisfied from the definition of $x(b',b,y(h))$. The second condition is satisfied if $x(b'_2,b,y(h)) - x(b'_1,b,y(h))$ is increasing in b for $b'_2 > b'_1$, provided that b'_1 and b'_2 are feasible. From the definition of $x(b',b,y(h))$, it follows that $x(b'_2,b,y(h)) - x(b'_1,b,y(h))$ is independent of, and hence increasing in, b. This resembles part (a) of Lemma 2 in the online appendix of Gordon and Qiu (2018). Lastly, to check whether $\mathbf{W}(b',h)$ is increasing in b', note that the continuation value depends on b' only through the value of repayment. The repayment value must be increasing in assets since the choice set expands in assets. Therefore, also the third condition holds, and we can conclude that the savings policy conditional on the productivity level is monotone.

3.D Solving the individual planning problem

To solve the household problem, we use the fact that at an interior solution, the following first-order condition must hold:

$$u'(x) = \frac{\beta}{q(b',h) + b'\frac{\partial}{\partial b'}q(b',h)}\mathbb{E}\left[\frac{u'(x')}{\pi}\mathbb{P}(D'=0|b',h')\middle| h\right] \tag{3.37}$$

which follows from the expression for the expected value of $V(b,h)$ with respect to shocks ν and the envelope condition. Note here that we divide by steady-state inflation π, which of course becomes π' when we linearize the model. The typical Laffer-Curve relationship between $q(b',h)b'$ and b' implies that it is never optimal to choose asset levels which are located on the downward sloping part of $q(b',h)b'$. In other words, only asset levels for which the first denominator on the RHS is positive are candidates for an optimal solution. Our algorithm then proceeds as follows. We first define an exogenous grid for future asset levels b' and current productivity levels h that we keep during iterations. Then, we guess initial value functions, marginal utilities, a price function, and its derivative. At this point, all terms on the RHS of the Euler equation are available at all grid points. Next, we apply the inverse marginal utility function to the RHS to calculate endogenous grid points for composite consumption $\tilde{x}$ for each productivity level and all b' for which $\frac{\partial}{\partial b'}q(b',h)b' > 0$. We also save the location of the smallest grid point for which the last condition is satisfied, which we denote by $b'_L(h)$. Using the budget constraint, we can back out the corresponding endogenous asset grid points $\tilde{b}$. Up to this point, nothing is different from the original algorithm in Carroll (2006). Now, however, we have to check whether these solutions to the Euler equation are in fact optimal. As the problem is not globally concave, the first-order condition above is only necessary, but not sufficient. Since we know that the optimal policy is monotone in beginning-of-period assets, the algorithm checks whether the endogenous grid points are increasing. If they are, we keep all of them. If not, we run a refinement step as in Fella (2014). Specifically, given the endogenous grid points as current assets, we run an on-grid maximization step to determine if the corresponding exogenous grid points are truly maximizers, and discard points where this is not the case. One issue that remains is that the number of endogenous points may be smaller than that of the exogenous points, since we did not necessarily invert the Euler equation at every asset grid point. To find a policy also at these points, we extrapolate the savings function from $\tilde{b}$ to b' at these points, but replace all extrapolated values which lie on the downward sloping part of the Laffer-Curve by $b'_L(h)$, because it is never optimal to choose asset levels below this point.

With the savings policy in hand, all remaining objects can be calculated to produce new values for the value functions and marginal utilities. Generally, it can happen that some choices are in fact no feasible. In this case, we set the repayment value to a large negative number (-9999) and marginal utility to a large positive number (9999) at these points. Then, the default probabilities are updated, which gives a new price schedule. To determine the derivative of the product $q(b', h)b'$, we make use of the fact that it can also be expressed in closed form as

$$\frac{\partial}{\partial b'} q(b', h)b' = q(b', h) + \frac{b'}{R} \mathbb{E}\left[\frac{1}{\sigma_\nu} \mathbb{P}(d' = 1|b', h')(1 - \mathbb{P}(d' = 1|b', h')) \frac{u'(x')}{\pi} \middle| h \right] \tag{3.38}$$

This expression follows from the expression for the default probabilities in the main text and the envelope condition. At this stage, we have solved the planning problem given bond prices, and determined new bond prices given this solution. We then repeat this process until all functions converge.

Practical issues: When solving the model, we found that two issues where important to make the algorithm perform better numerically, besides finding a suitable size of the taste shock variance. The first concerns the calculation of the expected value of $V(b, h)$ with respect to the taste shocks, and default probabilities. Namely, instead of using the formulas in the main text, in the codes we use the following equivalent expressions:

$$\begin{aligned} \mathbb{E}_\nu \left[V(b, h) \right] &= M + \sigma_\nu \log \left(\exp((W^D(h) - M)/\sigma_\nu) + \exp((W^R(b, h) - M)/\sigma_\nu) \right) \\ &\equiv \mathbf{EV}(b, h) \end{aligned} \tag{3.39}$$

for the expected value over taste shocks, where $M = \max(W^R(b, h), W^D(h))$, and

$$\mathbb{P}(d = 1|b, h) = \exp((W^D(h) - \mathbf{EV}(b, h))/\sigma_\nu) \tag{3.40}$$

for default probabilities. These formulas avoid taking exponentials of large negative numbers.

The second issue concerns the Laffer-Curves. Once we have found the savings policies, to find consumption we need to evaluate the product $q(b', h)b'$ at these policies, which are typically not on the grid. When we linearly interpolated the bond prices and multiplied by the policy, we found that steady-state profits in the banking sector were quiet large. This comes from the fact that in (3.9) we canceled b' on both sides, but linear interpolation between grid points of the bond prices assigns a wrong value to the bond price in between these points, which leads to biased results. We found that computing $q(b', h)b'$ and interpolating this product lead to steady state bank profits which were negligible.

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