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Crisis? What Crisis? Currency vs. Banking in the Financial Crisis of 1931

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Abstract

This paper examines the role of currency and banking in the German financial crisis of 1931 for both Germany and the U.S. We specify a structural dynamic factor model to identify financial and monetary factors separately for each of the two economies. We find that monetary transmission through the Gold Standard played only a minor role in causing and propagating the crisis, while financial distress was important. We also find evidence of crisis propagation from Germany to the U.S. via the banking channel. Banking distress in both economies was apparently not endogenous to monetary policy. Results confirm Bernanke's (1983) conjecture that an independent, non-monetary financial channel of crisis propagation was operative in the Great Depression.

JEL: N12, N13, E37, E47, C53
Keywords: Great Depression, 1931 financial crisis, international business cycle transmission, Bayesian factor analysis, currency, banking

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1 Introduction

Between 1929 and 1932, national output in the U.S. and Germany declined in unison, earlier and more strongly than in most other industrialized nations (see the data in Barro and Ursúa, 2008). The two economies were heavily exposed to each other, both through financial markets and the Gold Standard. German commercial debt owed directly and indirectly to the U.S. exceeded 10% of U.S. GDP in 1931. German reparations, owed indirectly to the U.S. through inter-allied loans from WW1 for which they served as collateral, again exceeded 10% of U.S. 1931 GDP. Both classes of debt were lost almost entirely between 1931 and 1933 (Schuker, 1988). The trigger event for this was the Austro-German financial crisis of July 1931. In a matter of days, it led to the nationalization of Germany's five largest banks, the suspension of gold convertibility, the introduction of capital controls, and a moratorium on reparations (see James, 1986, for an account of events).

Schnabel (2004) highlighted the vulnerability of German banks as a main cause of the 1931 crisis, identifying lack of equity and high exposure to short-term foreign credit as key factors. The weak position of Germany's banks had been inherited from the stabilization after the hyperinflation of 1923, which was strongly based on U.S. credit.

The 1931 financial crisis was also the first major crisis of the interwar Gold Standard, and effectively marked the beginning of its breakdown. Doubts about the credibility of Germany's commitment to the Gold Standard, as well as its ability to defend its currency, were emphasized by Eichengreen (1992) and Temin (1989).

Moreover, the financial crisis of 1931 was a foreign debt and reparation crisis. Large foreign borrowing under the favorable terms of the Dawes Plan between 1924 and 1929 had diluted the value of reparation claims. Stricter terms for reparation payments under the Young Plan helped to dry out further lending to Germany and led to a policy of fiscal austerity (Ritschl, 2002b). Dwindling domestic support for this policy in early 1931 triggered doubts about Germany's willingness and ability to pay further reparations, which contributed to the outbreak of the crisis.

Scholars have long emphasized the fact that both Germany's financial system and its foreign public debt were mainly underwritten by the U.S., see Kindleberger (1973) and in particular, Schuker (1988). This would make spillover effects of Germany's crisis on the U.S. seem plausible. Following Friedman and Schwartz (1963), historians have seen the financial crises of 1931 as one link in a chain of events that helped to turn the U.S. recession after 1929 into a catastrophic recession (see Temin, 1989). Bernanke (1983) argued that these financial crises operated as an independent, non-monetary channel of crisis transmission and propagation during the Great Depression.

The present paper is about identifying this financial channel and assessing its importance in aggravating the Great Depression in 1931. We employ dynamic factor

analysis (DFA) to aggregate the information in a large number of financial, monetary, and real time series from both the U.S. and Germany. Our choice of the U.S./Germany comparison is motivated both by the dominant role of the U.S. as the anchor of the interwar gold standard and the high mutual financial exposure of the U.S. and Germany. We provide structure to the factor model by exclusion restrictions on the factor loadings. For each country, we specify a currency component, a banking factor, and a real component separately. The first is designed to capture monetary transmission channels under the Gold Standard, which would be in line with more traditional interpretations of the 1931 crisis as first- or second-generation currency crisis (as in Eichengreen, 1992 or Temin, 2008). The banking component is designed to a measure of financial distress, reflecting views of the German 1931 crisis as a banking crisis by Schumpeter (1939), Born (1967) and James (1986), or more recently, as a third generation twin crisis (see Kaminsky and Reinhart, 1998, and the ensuing literature) by Schnabel (2004) and Adalet (2005).

The presence of identified common components in both countries allows us to examine their dynamic relationships both domestically and internationally. We do this obtaining impulse response functions from the factors under weak identifying restrictions. We also assess the information content of the individual factors by measuring their contribution to the forecasting power of the dynamic factor model. We do this at several critical junctures before and during the crisis, trying to obtain a pattern causality and propagation.

The idea that transmission of the 1931 financial crisis to the U.S. was important was emphasized by James (2001, 2009). Coincident with the German banking crisis, Richardson and van Horn (2008) find a strong increase in financial distress at New York banks. Accominotti (2009) examined bank balance sheets from London and found that the German banking crisis was instrumental in weakening the Sterling and pushing Britain off the Gold Standard. Mouré (2002) argued that after the end of Germany's reparations in August 1932, France's default on her portion of the inter-allied debt in 1932, along with her gold withdrawals, seriously worsened the credit crunch in the U.S. (see also Eichengreen and Flandreau, 2008).

Our results indicate that both monetary and financial transmission mechanisms were active during the slump. However, financial factors constitute by far the dominant channel of internaitonal crisis propagation, while monetary forces played only a moderate role (using a DSGE model Cole, Ohanian, and Leung, 2005, obtain related results). This also holds domestically for both economies, which is consistent with evidence from a FAVAR model for the U.S. in Amir Ahmadi and Ritschl (2009). We also find that contrary to expectation, crisis transmission from the U.S. to Germany was comparatively minor. In contrast, we obtain evidence of marked feedback effects from Germany on the U.S., transmitted mainly through the financial stress components.

These feedback effects became pronounced around the German crisis of July 1931. We find strong predictive power of Germany's financial factor for the U.S. economy, indicating a strong systemic component of the July 1931 crisis. We also find evidence that shock transmission to the U.S. after the crisis is stronger than before.

Our results relate closely to research in recent years about foreign debt crises

and their international spillovers. Calvo, Leiderman, and Reinhart (1996) have identified large output effects of such crises in the defaulting countries as well as marked spillover effects. Calvo, Izquierdo, and Talvi (2006) have argued that the U.S. depression of 1929 to 1933 and the subsequent recovery to 1937 bear a lot of resemblance to foreign-debt-related recessions. With due caution, our results on the transatlantic spillover of Germany's financial crisis can be viewed as complementary to and consistent with this interpretation.

To analyze the issue econometrically, we chose an approach that allows for sufficiently rich dynamics while capturing information from a large number of time series. Vector autoregression (VAR) analysis alone would not be the adequate tool because of its limitation to hardly more than a few time series. To exploit the information imbedded in many disaggregate time series and avoid the curse of dimensionality, we rely on a dynamic version of factor analysis as e.g. in Forni, Hallin, Lippi, and Reichlin (2000) or Stock and Watson (2002a,b). As indicated above, we combine the dynamic factor model with vector autoregressions to analyze the interdependencies between the estimated latent factors, following the factor augmented vector autoregression (FAVAR) approach by Bernanke, Boivin, and Eliasz (2005). Our version of the FAVAR model identifies the factors by exclusion restrictions, thus giving them a structural interpretation (as in Kose, Otrok, and Whiteman, 2003). However, we do not attempt to identify monetary policy instruments, as the focus of our interest is less on policy impulses but rather on the channels of transmission themselves.

Our approach to the dynamic factor models is a Bayesian one. We employ Monte Carlo Markov chain (MCMC) techniques to infer the posterior distributions. Our choice of a Bayesian framework is motivated by pragmatic considerations regarding computational convenience, following the lead of Otrok and Whiteman (1998) and Kim and Nelson (1998). As is implicit in the MCMC methodology, our estimates are quite robust to changes in the prior; hence our choice of the Bayesian framework can be regarded as a matter of computational convenience. The Bayesian approach also suggests itself from our choice of a structural factor model, as Bayesian numerical techniques are particularly robust in the presence of identifying exclusions restrictions.

Business cycle transmission with recent international data has been analyzed by structural VARs e.g. in Stock and Watson (2005) and by dynamic factor models in Eickmeier (2007). To our knowledge, the present paper is the first study applying modern time series methodology to the international transmission of the interwar Great Depression. Due to the limitations that existed so far in extending VARs to panel data, existing econometric work on the international Great Depression, as in Bernanke and James (1991) and Bernanke and Carey (1996), was confined to cross section methods.

We structure the evidence by grouping the national time series into nominal and real series and extracting identified factors specific to these groups under exclusion restrictions. We find that the real factors we construct from the data coincide well with traditional business cycle dating schemes and historical national accounts for the respective countries. This is well in line with the results of Stock and Watson (1998) on a factor approach towards business cycle dating. We group the nominal series further by subdividing them into general monetary indicators on the one hand

and bank specific indicators on the other. The factors we extract from these series again seem to replicate the historical evidence well.

The rest of this paper is structured as follows. The next section characterizes the dynamic factor model we employ. Section 3 provides the data. Section 4 obtains the factors and evaluates the relative importance of currency and banking in the German crisis. Section 5 concludes.

2 A Structural DFA Model

The dynamic factor approach is to to assemble more information than could be processed by a standard VAR analysis, the workhorse model of empirical macroeconomic analysis. We follow recent developments in dynamic factor analysis that have augmented VARs with information gathered from a large cross section of time series. The idea is to aggregate the common components of large time series panels into synthetic series or factors, which are then used as inputs into a standard VAR. For each of the two economies in our dataset, we restrict the factor loadings to specific subsets of the series, monetary, financial, and real.

The data panel Y_t , spanning a cross section of N series and an observation period of length T, is described by the following equation:

$$Y_t = C + \Lambda f_t + U_t \tag{1}$$

where f_t is a $K \times 1$ vector containing the latent factors, U_t is a $N \times 1$ vector of variable-specific idiosyncratic components, C is an $N \times 1$ vector of constant terms and Λ is the $N \times K$ coefficient matrix linking the K common factors to the i-th variable. More precisely, the Λ matrix controls for the structural interpretation of the factors, where each factor can be loaded on a subset of the data by imposing zero restrictions. In this context, we define

$$\Lambda = \begin{bmatrix} \Lambda^{US} & 0 \\ 0 & \Lambda^{D} \end{bmatrix}$$

where

$$\Lambda^{US} = \begin{bmatrix} \Lambda^{real} & 0 & 0 \\ 0 & \Lambda^{monetary} & 0 \\ 0 & 0 & \Lambda^{financial} \end{bmatrix}$$

and

$$\Lambda^{D} = \begin{bmatrix} \Lambda^{real} & 0 & 0 \\ 0 & \Lambda^{monetary} & 0 \\ 0 & 0 & \Lambda^{financial} \end{bmatrix}$$

The law of motion for the factors, which is in VAR form, is defined as:

$$f_t = \phi_1 f_{t-1} + \cdots + \phi_q f_{t-q} + v_t, \tag{2}$$

with $v_t \sim \mathcal{N}(0, \Sigma)$. The idiosyncratic components U_t are assumed to follow an AR(p) process:

$$U_t = \Theta_1 U_{t-1} + \ldots + \Theta_p U_{t-p} + \chi_t \tag{3}$$

where $\Theta_1, \ldots, \Theta_p$ are $N \times N$ diagonal matrices and $\chi_t \sim \mathcal{N}(0_{N\times 1}, \Omega_\chi)$ with

$$\Omega_\chi = \begin{bmatrix} \sigma^2_{1,\chi} & 0 & \cdots & 0 \\ 0 & \sigma^2_{2,\chi} & \vdots & \vdots \\ \vdots & \cdots & \ddots & 0 \\ 0 & \cdots & 0 & \sigma^2_{N,\chi} \end{bmatrix}$$

To ease the computational burden we *quasi difference* equation (1). Accordingly we multiply equation (1) by $(\mathcal{I} - \Theta(L))$, where $\Theta(L) = \Theta_1 + \cdots + \Theta_p$ and $\mathcal{I}$ is the identity matrix, which leads to the following expression:

$$Y^*_t = C^* + \Lambda^* f_t + \chi_t, \tag{4}$$

where $Y^*_t = (\mathcal{I} - \Theta(L))Y_t$, $\Lambda^* = (\mathcal{I} - \Theta(L))\Lambda$ and $C^* = (\mathcal{I} - \Theta(L))C$.

Prior Specification

For the AR-Parameters of the idiosyncratic components $\Theta_1, \Theta_2, \ldots, \Theta_p$ we specified the following prior:

$$\theta^{prior} \sim \mathcal{N}(\underline{\theta}, \underline{V}_\theta)$$

where $\underline{\theta} = 0_{p\times 1}$ and where

$$\left[\underline{V}_\theta\right] = \tau_1 \begin{bmatrix} 1 & 0 & \cdots & 0 \\ 0 & \frac{1}{2} & \vdots & \vdots \\ \vdots & \cdots & \ddots & 0 \\ 0 & \cdots & 0 & \frac{1}{p} \end{bmatrix}$$

We choose $\tau_1 = 0.2$. The shrinkage prior we specified implies that we punish more distant lags. This is applied by subsequently decreasing the uncertainty about the mean prior belief that the parameters are zero for increasing lag values.
For each of the factor loadings we specified the following prior:

$$\lambda^{prior} \sim \mathcal{N}(\underline{\lambda}, \underline{V}_\lambda)$$

where $\underline{\lambda} = 0$ and $\underline{V}_\lambda = 100$. For each of the variances of the disturbances in χ_t we specified the following prior:

$$\sigma^{prior}_\chi \sim \mathcal{IG}\left(\frac{\alpha_\chi}{2}, \frac{\delta_\chi}{2}\right)$$

where we choose $\alpha_\chi = 6$ and $\delta_\chi = 0.001$, which implies a fairly loose prior. $\mathcal{IG}$ denotes the inverted gamma distribution.

For the parameters of the VAR equation (2) we follow Bernanke, Boivin, and Eliasz (2005) and impose the Kadiyala and Karlsson (1997) Minnesota-type prior on the

VAR parameters. Then, the prior distribution of the covariance matrix Σ and the VAR parameters Φ can be expressed by:

$$\Sigma_{prior} \sim \mathcal{IW}(\underline{\Sigma}, K+2),$$

with $\mathcal{IW}$ representing the inverse Wishart distribution and

$$vec(\Phi_{prior}) \sim \mathcal{N}(0, \Sigma_{prior} \otimes \underline{G}),$$

where $\underline{G}$ imposes less weight on more distant lags .

2.1 Estimation

Estimation of the model is via the Gibbs sampler. The principal idea of this algorithm is to break the joint distribution of the model parameters into the conditional distributions and to proceed by iterating over the conditional distributions. As a first step, we start by drawing the parameter block $\Xi = [\Lambda, \Theta_1, \ldots, \Theta_p, \Phi, \Omega_\chi \Sigma]$ and take values for the factors as given. In the next step we use the obtained draws and calculate the factors conditional on the realizations of the previous block. These values of the first Gibbs Sampling step are now used to compute the next step by iterating through the blocks just mentioned. Iterating over sufficiently many steps, the simulated frequency distribution converges to the joint distribution at an exponential rate.[1]. To ensure that the dynamic factor model is uniquely identified, the upper $K \times K$ block of the factor loadings matrix is set to the identity matrix[2] where each diagonal element corresponds to one of the structural factors.

3 Data

Data are at a monthly frequency from September 1925 to November 1932. The U.S. series are taken from the NBER's macroeconomic history database, while the German data we take from Wagemann (1935). The U.S. data include, among others, bank debits, deposits, discount rates, steel production, machinery prices, orders of machinery, as well as an index of industrial production and trade. The German series are, among others, short term deposits, wholesale and consumer price indices, currency in circulation, discount rates, domestic orders of machinery, steel production, industrial production, and employment in the metal trades. All data except for the interest rates were standardized and transformed into first differences. For a more detailed description of the dataset see Appendix B.

4 Results

For the empirical results we choose the lag lengths $p = 1, q = 7$. We cycled through 30,000 Gibbs iterations. To avoid that our results are driven by the starting values we

[1] See Geman and Geman (1984) A more detailed description of the estimation procedure is provided in Appendix A.

[2] This is again similar to Bernanke, Boivin, and Eliasz (2005).

discard the first 10,000 draws of the chain as burn-in. We ensured global convergence by restarting the algorithm several times over, each time using different starting values drawn from an overdispersed distribution. Results obtained were very similar. In each case, the sampler reached convergence already after a few thousand draws.

4.1 Real and Nominal Factors

To add structure to the factor approach, we restrict the data space on which factors are allowed to load. For both the U.S. and Germany, we identify three factors, one of them real, the other two nominal. The first factor is designed to capture real activity in the respective national economies. The two nominal factors load on a number of currency and banking series, respectively.

(Figure 1 about here)

The real factor for the U.S. loads on output data for investment goods, as well as a contemporary index of output in manufacturing and trade. The result is shown in Figure 1(a). This factor is essentially a reflection of traditional business cycle chronologies, and is highly correlated with the most commonly used indices of industrial production. We found the result to be very robust to changes in the specification of the time series included. We also notice a very good fit with a broadly based factor of economic activity calculated in Ritschl, Sarferaz, and Uebele (2008). Our results confirm the observation by Stock and Watson (1998) that one-factor models describe the real state of the economy quite well.

The monetary factor for the U.S. in Figure 1(c) loads on different short-term interest rates. By construction, this factor closely mirrors the increase in short term interest rates through late 1929, followed by a sharp decline to early 1931. A pronounced upward shock becomes visible in mid-1931, right around the time of Germany's 1931 crisis.

The U.S. banking factor in Figure 1(e) is based on the commonly used banking statistics from the NBER database. It shows continuing expansion through the 1920s, and reaches its peak with the October 1929 crash. The banking panic of December 1930 is also visible. Again, there is an additional downward shock in mid-1931, right after the German crisis.

Figure 1(b) shows the German factor of real activity: fast recovery from a recession in 1925/6 is followed by a marked slowdown in 1927. Real activity peaks in the summer of 1929, and is already in decline at the time of the New York stock market crash. A beginning recovery in the first half of 1931 is suddenly chocked off by a strong downward shock at the time of the German crisis. After a double dip in summer 1932, recovery set in and was well under way before early 1933, when the Nazis got to power. All this is in line with conventional wisdom (see Ritschl, 2002a for a discussion).

The German currency factor in Figure 1(d) is again largely composed of interest rates. It peaks in mid-1929 and then falls rapidly to reach its trough in mid-1930. An upward jump is visible in September 1930, after a national election that sharply increased the Nazi and communist votes. There is some slight improvement before

the German crisis of mid-1931 and a huge shock afterwards. Interest rates came down markedly during 1932, leveling out towards the end of 1932.

The banking factor in Figure 1(f), loading on the banking series in our dataset, is rather similar to series generated by Schnabel (2004) and Adalet (2005). It shows almost continuous improvement to March 1929, when a first setback occurred, coincident with the first Young Plan crisis (see James, 1985). Recovery to early 1930 was followed by a second setback, coincident with the adoption of the Young Plan, Schacht's resignation from the Reichsbank presidency, and the downfall of the last parliamentary government. After that, the banking factor begins a precipitous decline, which develops into a collapse at the time of the mid-1931 crisis. There is no recovery until early 1933. Germany's two nominal factors thus both show a major, sudden decline in mid-1931. Eyeballing the evidence from the factors, one may conclude that both a currency and a banking crisis were at work.

Drawing the evidence from this section together, a common salient feature of the factors, and thus of the common underlying dynamics of our time series, is the marked deterioration in mid-1931, at the time of the German crisis. This effect is not limited to the German data, and is indeed visible also in the factors we extracted from the U.S. series. The next section will trace the phenomenon further, employing impulse-response analysis of a structural FAVAR.

4.2 Currency vs. Banking: the Transmission of Shocks

This section relates the above factors to each other in a VAR analysis. As the factors have a structural interpretation, the dynamic relationships between these factors can be given a structural interpretation as well. This section analyzes the transmission of surprise shocks across the two economies using impulse response functions. Our interest focuses on the relative importance of monetary shocks, transmitted through the Gold Standard mechanism, and of shocks to the banking system, transmitted through the mutual exposure of the two countries' banking systems to each other.

We orthogonalize the shocks using mostly the temporal Cholesky decomposition. Our principal identification strategy is to assume that the U.S. factors do not react simultaneously to international conditions, while the German ones do: U.S. real activity is assumed endogenous to U.S. monetary and banking conditions only. German currency conditions are assumed endogenous to U.S. factors but exogenous to German banking conditions. We furthermore assume that German real were endogenous to all other factors.

The only exception to this identification strategy is the propagation of shocks to the U.S. interest and banking factors, for which we adopt the agnostic sign restriction approach of Uhlig (2005). The idea is to focus only on those results that yield plausible impulse responses for the nominal side of the economy, while being agnostic with regard to the response of real activity in the economy. Uhlig (2005) suggested this approach as an alternative to the recursive Cholesky identification in order to avoid sign puzzles in the response of nominal series to a monetary shock at short horizons. Such sign puzzles would abound in impulse responses obtained via the Cholesky decomposition from U.S. interwar data, which makes the use of an alternative approach compelling.

To identify the responses to nominal shocks, we present results for two alternative sets of sign restrictions. In a baseline identification, we restrict the responses of both the U.S. monetary and the U.S. banking factor to a nominal shock to be negative for six months. No sign restriction operates on the responses of real activity in both countries to a nominal shock. We also experiment with a departure from Uhlig's (2005) agnostic approach toward the response of real activity and employ a stronger identification, restricting the response of U.S. output to be negative as well.

To account for the potential effects of the German crisis of July 1931, we also run the FAVAR analysis of this section separately for a truncated observation period from 1925 to May 1931, cutting off just before the onset of the financial crisis. Comparison of the impulse response functions from the full and truncated sample allows us to draw additional conclusions about the possible impact of the 1931 crisis.

4.2.1 Full Observation Period

Figure 2 shows the impulse response functions and the error bands for adverse shocks to U.S. real activity. Such shocks tended to be quite persistent. They were transmitted to the U.S. monetary factor, which exhibits a marked downward response of interest rates. Strong adverse effects on U.S. banking conditions existed as well. On average, around 40% of the forecast error variance in the U.S. banking factor is explained by real shocks, albeit with substantial error margins. The German economy shows similar responses to real shocks on U.S. economy, albeit in weakened form.

(Figures 2 and 3 about here)

To identify the effects of nominal shocks to the U.S. economy, we proceed in two steps. Figure 4 shows the responses to an adverse nominal shock, were the responses of both the monetary and the banking factors for the U.S. are restricted to be negative for six months. Under this baseline, the responses of Germany's nominal factors over the same horizon are negative as well. This seems like a desirable property: an identified nominal shock to the U.S. operates like a global nominal shock, the two are observationally equivalent.

The real factor in both economies also respond in almost identical fashion, however with less desirable properties. The median response of U.S. real activity over a six-month horizon is just negative, indicating that almost 50% of the draws are positive. The response of German real activity to a nominal shock is equally diffuse, again with almost half of the probability mass in the positive orthant. For both countries, the forecast error variance in real activity explained by the nominal shock is minimal, averaging less than 10%, see Figure 5.

(Figures 4 and 5 about here)

To achieve sharper results for the real responses to nominal shocks, we depart for a moment from Uhlig's (2005) agnostic stance on output responses and force the response of U.S. real activity to be negative for six months after a nominal shock. This additional constraint allows us to identify shocks to monetary conditions and

to the banking factor separately.

Figure 6 shows the responses to tightening conditions in the U.S. money market (although not necessarily to monetary policy itself). By construction, the response of U.S. real activity is now negative for six months. This sign restriction is a binding constraint. In its absence, the response of the U.S. real factor to a monetary shock would have been positive throughout. On average, U.S. banking responds negatively for most horizons, although large parts of the probability mass indicate positive responses. As suspected by Bernanke (1983), monetary factors have only limited explanatory power for financial conditions: hardly more than 10% of the forecast error variation in the U.S. banking factor are explained by the U.S. interest factor. This result was very robust under a variety of different specifications of both the monetary and the ban king factors. The sign restriction on the monetary factor itself is again binding: as soon as the constraint is lifted, the response turns into negative territory. The responses of the German factors are similar to their U.S. counterparts but on the whole appear more diffuse.

(Figures 6 and 8 about here)
(Figures 7 and 9 about here)

The forecast error decompositions in Figure 7 suggest a share of 10-20% for nominal tightening in explaining the variance of U.S. real activity. This appears to confirm results of Sims (1999) in a longitudinal study of U.S. monetary policy in the 20th century, as well as of Amir Ahmadi and Ritschl (2009) from a FAVAR model for U.S. monetary policy during the Great Depression.

Figure 8 shows the responses to tightening financial conditions. Again obtained under sign restrictions, the shocks are quite persistent and also translate into persistent real effects. However, the response of real activity in the U.S. is negative for the first six months by construction. After that, it remains negative on average, but draws with positive responses do occur, indicating that the restriction is binding. Lifting the constraint, the responses would be positive throughout. As before, the German responses are structurally similar but more diffuse.

The forecast error decompositions in Figure 9 suggest that about 15 % of the variation in the real factor can be explained by shocks to financial conditions, which is slightly higher than for monetary shocks.

(Figures 10 and 12 about here)
(Figures 11 and 13 about here)

Next we look at the effects of shocks to the German factors. As would be expected, a shock to real activity in Germany (see Figure 10) is persistent domestically but has no discernible effect on the U.S. economy. Shocks to German money market conditions, shown in Figure 12, propagate through the German economy without sign puzzles and have real effects. However, their contribution to the forecast error variance of the German real factor is low (see Figure 13).

Monetary market tightening in Germany has near-significant effects on real conditions in the U.S., yet their contribution to forecast error variance is negligible

(see Figure 13). The effect of nominal tightening in Germany on U.S. monetary conditions is briefly negative and significant but then turns into positive, however without being significant. There is also a negative but insignificant effect on the U.S. banking environment. Both effects would be consistent with the classical gold standard mechanism, however in a slightly non-standard way: it almost looks as if the U.S. played the role of a monetary shock absorber for the international gold standard, much like the Bank of England in the pre-World War I years.

The same direction of causality becomes visible for adverse shocks to the German banking factor. Figure 14 shows persistent and significant effects on U.S. real activity as well as on U.S. banking conditions, while the effect on the U.S. interest factor is hump-shaped and changes signs. This effect of German banking conditions on U.S. conditions has hardly been studied so far; we found it to be robust under a large variety of alternative specifications. A look at the variance decompositions in Figure 15 shows a high contribution of Germany's banking factors to the forecast error variance of the German real factor. With a delay of about ten months, marked effects also build up on the variance of the U.S. real and banking factors.

(Figures 14 and 15 about here)

This result would lend support to the hypothesis of James (2001) that the deepening of the U.S. recession in 1931 was at least partly triggered by the international repercussions of the 1931 crisis in Austria and Germany. The variance decompositions in Figure 15 show that after two years, the cumulative effects of shocks to Germany's banking conditions on the U.S. real factor are markedly higher than for the U.S. monetary and banking factors in Fig. 6 and 8 above.

4.2.2 Truncated Observation Period, 1925 to June 1931

To identify the contribution of the 1931 crisis to this surprising result, we truncate the observation period to end in May 1931. Figure 16 shows the responses to German currency shocks for this subperiod. A surprising countercyclical pattern emerges: adverse monetary shocks in Germany have mostly adverse effects on German real activity in banking, but significant, favorable effects on real conditions in the U.S. In contrast, in this truncated sample from before the crisis, all responses to an adverse shock to German banking (in Figure 18) have roughly the same characteristics as for the whole sample (in Figure 14) but are less significant and have less explanatory power for forecast error variance. Evidently, the German crisis of 1931 sharpens the results. The financial accelerator effects of Germany's 1931 crisis on the U.S. economy must have been considerable.

(Figures 16 and 17 about here)
(Figures 18 and 19 about here)

Drawing the results of this section together, our application of a dynamic factor model finds little evidence for the traditional view that U.S. monetary or banking problems were key in explaining the depression in either country. We find only scant

support for a transmission of the recession from the U.S. to Germany through either monetary of financial channels of transmission. We also notice that nominal shocks to the U.S. economy do not play a dominant role in explaining the variation of real activity.

Conversely, we do find significant effects of Germany's nominal shocks on real activity in the U.S. economy. Again the monetary channel is of relatively minor importance. Transmission through the banking channel, however, comes out as quantitatively important and highly persistent. The effects have not fully built up after 20 months, and would explain 30% in the variance of both U.S. real activity and the U.S. banking factor.

However, most of these effects did apparently not really materialize before the 1931 crisis. Truncating the observation period to end in May 1931, we find the responses to Germany's nominal conditions to be less pronounced and less significant.[3] This implies that transmission from Germany to the U.S. is strongest in the period after July 1931. We conclude that international spillovers from the German crisis of 1931 were a significant force in deepening the U.S. recession.

We also find that while nominal factors seem to have played a rather minor role in the U.S. recession, the overall role of nominal factors in the German recession seems somewhat stronger. Responses of German real activity to adverse shocks in German monetary and banking conditions are estimated precisely and without having to resort to sign restrictions. In the case of financial shocks, they are also quantitatively important, accounting for a third of the forecast error variance in German real activity. Again, the explanatory power of monetary shocks is much lower: the explained variation in German real activity is only about 10%.

The results so far imply that banking conditions played a dominant role in the German crisis of 1931. As a corollary, if there was a financial frictions channel of transatlantic business cycle transmission in the Great Depression, it originated in Germany rather than in the U.S., and affected both economies significantly. This is consistent with the claim by Harold James (2001) that the German banking crisis had major spillover effects on the international economy. It is also consistent with the claim of James (1986) and Schnabel (2004) that Germany's 1931 crisis was causally a banking crisis, while monetary transmission under the Gold Standard played only a secondary role.

4.3 Currency vs. Banking: the Systematic Effects

Thus far, attention has focused on the transmission of surprise shocks. In the following section, we examine possible systematic effects that may have been factored into expectations. Systematic components included in the agents' information set at time t would be reflected in the accuracy of forecasts made on the basis of that information set. In this section we obtain forecasts of real activity in Germany and the U.S., conditional on the information at critical junctures before and during the 1931 crisis. To evaluate the information content of the banking factor at any of these points in time, we obtain each forecast twice, once from a bivariate VAR including

[3]The more direct test of obtaining the results for the subperiod from June 1931 to March 1933 would not be feasible due to missing degrees of freedom in specifying the model

the banking sector, once from a univariate AR of the same lag length in the real activity factor alone.

4.3.1 Germany

Univariate forecasts for the German real factor from March and May 1931 predict recovery, extrapolating from the green shoots that had become visible in early 1931. The forecasts are quite imprecise, though, with widely diverging error bands. Only if the update from July 1931, after the crisis, is incorporated does the univariate forecast predict a further downturn.

(Figure 20 about here)

To evaluate the gain in forecasting power from the information content in the banking factor, we now add the German Banking series and perform bivariate conditional forecasts for the same three truncated samples.

(Figure 21 about here)

The forecast of the German real sector from March 1931 already predicts further downturn, although large parts of the probability mass are still predicting a further increase. The forecast for May 1931 is much more unequivocal about a further decrease. Comparing this forecast to the univariate forecast for May in Figure 20 above, the banking series turns out to be highly informative about a renewed downturn. German banking variables up to May 1931 clearly predict a major deterioration before the July 1931 crisis. If the update to the banking series for July is included, we obtain a full prediction of the decline in real activity through mid-1931 (in Figure 21(c)).

No comparable gain in predictive power is obtained if we include monetary instead of banking variables in the forecasts. Results in Figure 22 show little improvement over the univariate forecast of real activity in Figure 20 above.

(Figure 22 about here)

Only if the information from July 1931 is incorporated does the bivariate forecast including monetary information predict the further decline in activity correctly.

These results confirm the evidence from the impulse response analysis in the previous section: the domestic driving force behind Germany's 1931 crisis was the weakness of its banking system. The deterioration in banking conditions foreshadowed the July 1931 crisis, and indeed has considerably predictive power. By comparison, domestic monetary conditions play only a secondary role.

No predictive power for German real activity is gained from including U.S. rather than German monetary and banking data in the forecasts before the July 1931 crisis.

(Figure 24 about here)
(Figure 25 about here)

Indeed it is noteworthy how the inclusion of the U.S. monetary factor tends to buttress the prediction of a continuing upswing in Spring 1931. Even after the beginning of crisis in July, the forecasts conditional on U.S. data are more optimistic than the univariate forecast in Figure 20 above. According to these results, U.S. data are uninformative about the German financial crisis; there is no indication that the 1931 crisis was triggered by conditions in the U.S.

4.3.2 U.S.

In Figure 27 we show forecasts for the U.S. real sector as of March 1931. As can be seen, they predict rather a stagnation than a further deterioration of the U.S. economy.

(Figure 27 about here)

The noteworthy exception is the forecast including the monetary factor, which predicts fast recovery. This would indicate that money market conditions were not a constraining factor in the spring of 1931.

Figure 28 shows the forecasts from May 1931 on. The univariate forecast is now more pessimistic, and the bivariate forecast including banking conditions is even more so. These are clear signs of mounting banking distress in the U.S. before the July 1931 crisis. In contrast, the forecast including monetary factors is again pointing to an imminent recovery.

(Figure 28 about here)

The forecasts from July 1931 confirm this result. Again, the bivariate forecast including banking activity is more pessimistic than the univariate forecast. It is also closer to the actual trajectory of real activity after the crisis. The forecast including monetary conditions once again comes out as more optimistic, signaling an end to the recession and a return to recovery in 1932.

(Figure 29 about here)

Again we examine mechanisms of transatlantic business cycle transmission, this time tracking possible anticipation and contagion effects of the German financial crisis on the U.S. Figure 30 shows bivariate forecasts of U.S. real activity from May 1931 including the German banking and monetary factors, respectively.

(Figure 30 about here)

The bivariate forecast of U.S. real activity using the German banking factor up until May 1931 is as pessimistic as the forecast using the U.S. banking factor in Figure 28 above. This result implies that German banking conditions in May 1931 were informative about U.S. real activity. In contrast, the German currency factor adds no predictive power and essentially reproduces the univariate forecast of U.S.

real activity in Figure 28 above.

For July 1931, including data from immediately after the German financial crisis we obtain a very similar result.

(Figure 31 about here)

German banking data are again highly informative about U.S. real activity; they now actually predict the further downturn slightly better than the bivariate forecast using the U.S. banking factor in Figure 29 above. In contrast, using German monetary information again fails to predict U.S. real activity and signals a swift (though short-lived) recovery.

In sum, we find that banking conditions in both the U.S. and Germany have considerable predictive power for real activity in mid-1931, while monetary factors do not. However, U.S. banking conditions have very little predictive power for German real activity, while the German banking factor is highly informative about U.S. real activity. This evidence would be difficult to reconcile with an interpretation of the 1931 financial crisis as a primarily monetary phenomenon, or as contagion of distress originating in the U.S. banking system. It is consistent, however, with the interpretation that the German financial crisis of 1931 was primarily rooted in Germany's national banking system and had strong adverse effects on the U.S. economy as well.

5 Conclusion

This paper assessed the relative importance, both domestic and international, of Gold Standard transmission vs. banking channels in the origins and the propagation of the German financial crisis of 1931. To identify channels of crisis causation and propagation, we employed a structural dynamic factor model of the interactions between the U.S. and the German economy between 1925 and 1932. To this end we restricted the model to yield structural factors representing banking and monetary conditions the U.S. and the German separately. We also included one real factor for each of the two economies. Our real factors appear to trace established business cycle chronologies very well. Our nominal factors for Germany suggest that both monetary and banking conditions in Germany deteriorated severely and persistently in the 1931 crisis.

The first main result of this paper is that the overall transmission of nominal shocks from the U.S. to the German economy was insignificant and quantitatively negligible. This implies only weak support for the conventional wisdom that monetary and banking conditions in the U.S. transmitted the Great Depression to the rest of the world. In spite of our use of a broad database, we do not detect the U.S. causation of the international depression that has been taken for granted in much of the traditional literature. In contrast, we find remarkably high transmission of shocks in U.S. real activity to the German economy.

A second main result of this study is that in both countries, the influence of domestic monetary conditions on real activity was weak. Neither surprise effects nor any systematic effects appear to play a significant role. This result proved

robust under a large variety of different specifications we experimented with.

The third main result of this paper is that banking conditions constitute a important channel of domestic propagation and international transmission of the Great Depression, confirming the central claim of Bernanke (1983). We find that banking conditions cannot be explained by monetary conditions but themselves have marked real effects. The domestic financial channel comes out stronger in Germany but is also present in the U.S. International transmission through the financial channel was from Germany to the U.S., from the periphery to the core. This effect comes out stronger after the 1931 crisis. We have argued in this paper that the U.S. was strongly exposed to the German economy through credit, and indirectly through reparations that collateralized inter-Allied war credits. Germany's banking system suffered a meltdown in mid-1931, which made this vulnerability visible. In the process, the U.S. lost loans to Germany and Europe that equaled Germany's GDP in 1931, or roughly on quarter of U.S. GDP in the same year. The collapse of Germany's financial position in 1931 was a key event in turning the international recession into an unprecedented economic disaster.

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A Estimation

A.1 Estimating the Parameter Block

In this section we condition on the factors f_t. Because equation (1) is set of N independent regressions with autoregressive error terms it is possible to estimate Λ, $\Theta_1, \Theta_2, \ldots, \Theta_p$, Ω_χ and Ω_ϵ equation by equation.[4] We rewrite equation (3) as:

$$u_i = X_{i,u}\theta_i + \chi_i \tag{5}$$

where $u_i = [u_{i,p+1}\ u_{i,p+2}\ \ldots\ u_{i,T}]'$ is $T-p\times 1$, $\theta_i = [\theta_{i,1}\ \theta_{i,2}\ \ldots\ \theta_{i,p}]'$, is $p\times 1$ and $\chi_i = [\chi_{i,p+1}\ \chi_{i,p+2}\ \cdots\ \chi_{i,T}]'$ is $T-p\times 1$ and

$$X_{i,u} = \begin{bmatrix} u_{i,p} & u_{i,p-1} & \cdots & u_{i,1} \\ u_{i,p+1} & u_{i,p} & \cdots & u_{i,2} \\ \vdots & \vdots & \vdots & \vdots \\ u_{i,T-1} & u_{i,T-2} & \cdots & u_{i,T-p} \end{bmatrix}$$

which is a $T-p\times p$ for $i = 1, 2, \ldots, N$.

Combining the priors described in section 2 with the likelihood function we obtain the following posterior distributions.

The posterior of the AR-parameters of the idiosyncratic components is:

$$\theta_i \sim N(\overline{\theta_i}, \overline{V}_{i,\theta}) I_{S_\theta} \tag{6}$$

where

$$\overline{\theta_i} = \left(\underline{V_\theta}^{-1} + (\sigma^2_{i,\chi})^{-1} X'_{i,u} X_{i,u}\right)^{-1} \left(\underline{V_\theta}^{-1}\underline{\theta} + (\sigma^2_{i,\chi})^{-1} X'_{i,u} u_i\right)$$

and

$$\overline{V}_{i,\theta} = \left(\underline{V_\theta}^{-1} + (\sigma^2_{i,\chi})^{-1} X'_{i,u} X_{i,u}\right)^{-1}.$$

where I_{S_θ} is an indicator function enforcing stationarity.

The posterior of the variance of the idiosyncratic component $\sigma_{i,\chi}$ is:

$$\sigma_{i,\chi} \sim \mathcal{IG}\left(\frac{(T+\alpha_\chi)}{2}, \frac{((u_i - X_i\theta_i)'(u_i - X_i\theta_i) + \delta_\chi)}{2}\right) \tag{7}$$

To estimate the factor loadings we rewrite equation (1) as:

$$y^*_i = c^*_i + \lambda_i f^* + \chi \tag{8}$$

where $y^*_i = [(1-\theta(L)_i)y_{i,p+1}\ (1-\theta(L)_i)y_{i,p+2}\ldots\ (1-\theta(L)_i)y_{i,T}]'$ which is $T-p\times 1$, $c^*_i = c_i(1-\theta(L)_i)$ and $f^* = [(1-\theta(L)_i)f_{p+1}\ (1-\theta(L)_i)f_{p+2}\ \ldots\ (1-\theta(L)_i)f_T]'$, which $T-p\times 1$ with $\theta(L)_i = (\theta_{i,1} + \theta_{i,2} + \cdots + \theta_{i,p})$ for $i = 1, 2, \ldots, N$. Thus, the posterior for the factor loadings is:

$$\lambda_i \sim N(\overline{\lambda_i}, \overline{V}_{i,\lambda}) \tag{9}$$

[4] See also Chib (1993).

where

$$\overline{\lambda_i} = \left(\underline{V}_\lambda^{-1} + (\sigma_{i,\chi}^2)^{-1} f^{*\prime} f^*\right)^{-1} \left(\underline{V}_\lambda^{-1} \underline{\lambda} + (\sigma_{i,\chi}^2)^{-1} f^{*\prime} y_i^*\right)$$

and

$$\overline{V}_{i,\lambda} = \left(\underline{V}_\lambda^{-1} + (\sigma_{i,\chi}^2)^{-1} f^{*\prime} f^*\right)^{-1}.$$

To estimate the VAR parameters of the factors $\phi_1, \phi_2, \ldots, \phi_q$ we find it useful to rewrite equation (2) as:

$$f = X_f \phi + \nu \tag{10}$$

where $f = [f_{q+1} \ f_{q+2} \ \ldots \ f_T]'$ is $T - q \times K$, $\phi = [\phi_1 \ \phi_2 \ \ldots \ \phi_q]'$ is $Kq \times K$, $\nu = [\nu_{q+1} \ \nu_{q+2} \ \ldots \ \nu_T]'$ is $T - q \times K$ and

$$X_f = \begin{bmatrix} f_q & f_{q-1} & \cdots & f_1 \\ f_{q+1} & f_q & \cdots & f_2 \\ \vdots & \vdots & \vdots & \vdots \\ f_{T-1} & f_{T-2} & \cdots & f_{T-q} \end{bmatrix}$$

which is $T - q \times Kq$. Thus, the posterior of the VAR parameters can be drawn from the following distribution:

$$vec(\Phi) \sim \mathcal{N}(vec(\bar{\Phi}), \Sigma \otimes \overline{G}) I_{S_\Phi},$$

where $\bar{\Phi} \equiv \overline{G}(X_f{}' X_f)\hat{\Phi}$ and $\overline{G} = (\underline{G}^{-1} + X_f{}' X_f)^{-1}$. where I_{S_Φ} is an indicator function enforcing stationarity.

A.2 Estimating the Latent Factors

To estimate the common latent factor we condition on the parameters of the model.[5] Our observation equation in the following state-space system is:

$$Y_t^* = C^* + H F_t + \chi_t \tag{11}$$

where

$$H = [\Lambda \ -\Theta_1 \Lambda \ -\Theta_2 \Lambda \ \ldots \ \Theta_p \Lambda \ 0_{N \times K(q-p-1)}]$$

Our state equation is:

$$F_t = \Phi F_{t-1} + \widetilde{\nu}_t \tag{12}$$

where $F_t = [f_t, f_{t-1}, \ldots, f_{t-q+1}]'$ is $Kq \times 1$, which is denoted as the state vector, $\widetilde{\nu}_t = [\nu_t \ 0 \ \ldots \ 0]'$ is $Kq \times 1$ and

$$\Phi = \begin{bmatrix} \phi_1 & \phi_2 & \cdots & \phi_q \\ & \mathcal{I}_{K(q-1)} & & 0_{K(q-1) \times K} \end{bmatrix}$$

which is $Kq \times Kq$. For all empirical results shown below we use $q > p$.

[5] See also Kim and Nelson (1999)

To calculate the common factor we use the algorithm suggested by Carter and Kohn (1994) and Frühwirth-Schnatter (1994) . This procedure draws the vector $F = [F_1\ F_2\ \ldots\ F_T]$ from its joint distribution given by:

$$p(F|\Xi, Y) = p(F_T|\Xi, y_T) \prod_{t=1}^{T-1} p(F_t|F_{t+1}, \Xi, Y^t) \tag{13}$$

where $\Xi = [\Lambda, \Theta_1, \ldots, \Theta_p, \Phi, \Sigma, \Omega_c hi]$ and $Y^t = [Y_1\ Y_2\ \ldots\ Y_t]$. Because the error terms in equations (11) and (12) are Gaussian equation (13)can be rewritten as:

$$p(F|\Lambda, Y, \Xi) = \mathcal{N}(F_{T|T}, P_{T|T}) \prod_{t=1}^{T-1} \mathcal{N}(F_{t|t,F_{t+1}}, P_{t|t,F_{t+1}}) \tag{14}$$

with

$$\begin{aligned} F_{T|T} &= E(F_T|\Lambda, \Xi, Y) & (15)\\ P_{T|T} &= Cov(F_T|\Lambda, \Xi, Y) & (16) \end{aligned}$$

and

$$\begin{aligned} F_{t|t,F_{t+1}} &= E(F_t|F_{t+1}, \Lambda, \Xi, Y) & (17)\\ P_{t|t,F_{t+1}} &= Cov(F_t|F_{t+1}, \Lambda, \Xi, Y) & (18) \end{aligned}$$

We obtain $F_{T|T}$ and $P_{T|T}$ from the last step of the Kalman filter iteration and use them as the conditional mean and covariance matrix for the multivariate normal distribution $\mathcal{N}(F_{T|T}, P_{T|T})$ to draw F_T. To illustrate the Kalman Filter we work with the state-space system equations (11) and (12). We begin with the prediction steps:

$$\begin{aligned} F_{t|t-1} &= \Phi F_{t-1|t-1} & (19)\\ P_{t|t-1} &= \Phi P_{t-1|t-1}\Phi + Q & (20) \end{aligned}$$

where

$$Q = \begin{bmatrix} \Sigma & 0 & \cdots & 0 \\ 0 & 0 & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \cdots & 0 \end{bmatrix}$$

which is $Kq \times Kq$. To update these predictions we first have to derive the forecast error:

$$\kappa_t = Y_t^* - C^* - HF_{t|t-1} \tag{21}$$

its variance:

$$\Sigma = HP_{t|t-1}H' + \Omega_\chi \tag{22}$$

and the Kalman gain:

$$K_t = P_{t|t-1}H'\Sigma^{-1}. \tag{23}$$

Thus, the updating equations are:

$$\begin{aligned} F_{t|t} &= F_{t|t-1} + K_t\kappa_t, & (24)\\ P_{t|t} &= P_{t|t-1} + K_tHP_{t|t-1}, & (25) \end{aligned}$$

To obtain draws for $F_1, F_2, \ldots, F_{T-1}$ we sample from $\mathcal{N}(F_{t|t,F_{t+1}}, P_{t|t,F_{t+1}})$, using a backwards moving updating scheme, incorporating at time t information about F_t contained in period $t+1$. More precisely, we move backwards and generate F_t for $t = T-1, \ldots, p+1$ at each step while using information from the Kalman filter and F_{t+1} from the previous step. We do this until $p+1$ and calculate $f_1, f_2, \ldots, f_p$ in an one-step procedure.

The updating equations are:

$$F_{t|t,F_{t+1}} = F_{t|t} + P_{t|t}\Phi' P^{-1}_{t+1|t}(F_{t+1} - F_{t+1|t}) \tag{26}$$

and

$$P_{t|t,F_{t+1}} = P_{t|t} - P_{t|t}\Phi' P^{-1}_{t+1|t}\Phi P_{t|t} \tag{27}$$

B Data

	Series	Mnemonic
1	U.S. Steel Production	m01135a
2	U.S. Index of Industrial Production and Trade, Seasonally Adjusted	m12004c
3	U.S. Index of Orders for Machinery Tools and Forging Machinery	m06029
4	U.S. Index of Production Of Machinery, Seasonally Adjusted	m01277b
5	U.S. Index of Consumer Goods	m01056
6	U.S. Loans On Securities, Reporting Member Banks, Federal Reserve System	m14074
7	U.S. All Other Loans, Reporting Member Banks, Federal Reserve System	m14075a
8	U.S. Index of Deposit Activity	m12008b
9	U.S. Bank Debits	m12030
10	U.S. Clearings Index of Business	m12020b
11	U.S. Commercial Paper Rate	m13002
12	U.S. Discount Rates	m13009
13	U.S. Ninety Day Time-Money Rates On Stock Exchange Loans	m13003
14	German Orders of Machines	–
15	German Steel Production	–
16	German Industrial Production	–
17	German Employment in Metal Trade Sector	–
18	German Savings Deposits	–
19	German Demand Deposits	–
20	German Creditors	–
21	German Stocks of Bills of Exchange	–
22	German Debtors	–
23	German Discount Rates	–
24	German Private Discount Rates	–
25	German Warenwechsel	–

Source: German data are taken from Wagemann (1935). U.S. data are taken from the NBER macro history database, www.nber.org/databases/macrohistory/contents/.

C Figures

C.1 Latent Common Components

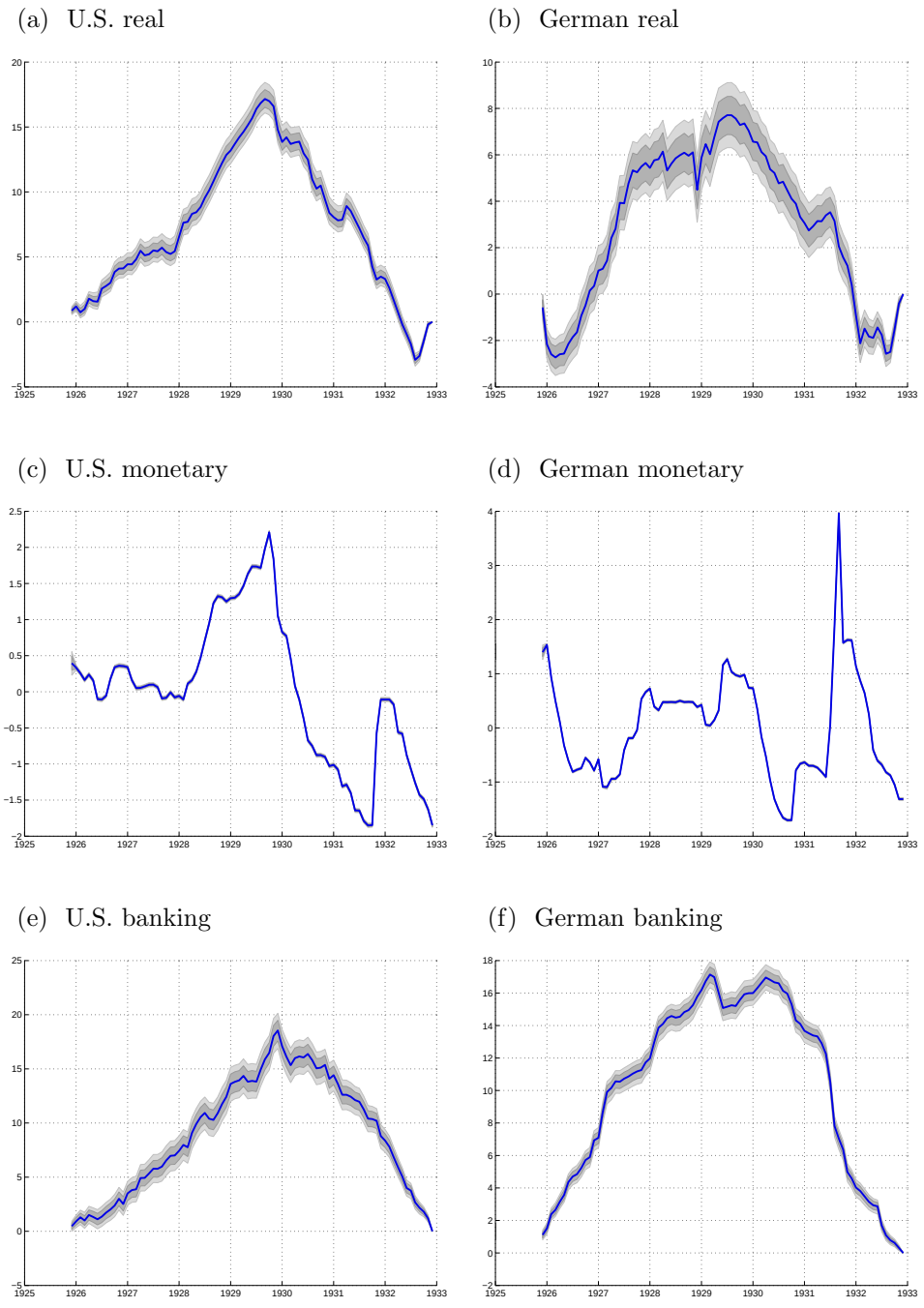


Figure 1: Latent common components for the U.S. and German real, monetary and financial variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

C.2 Impulse Response Analysis (1925:9–1932:11)

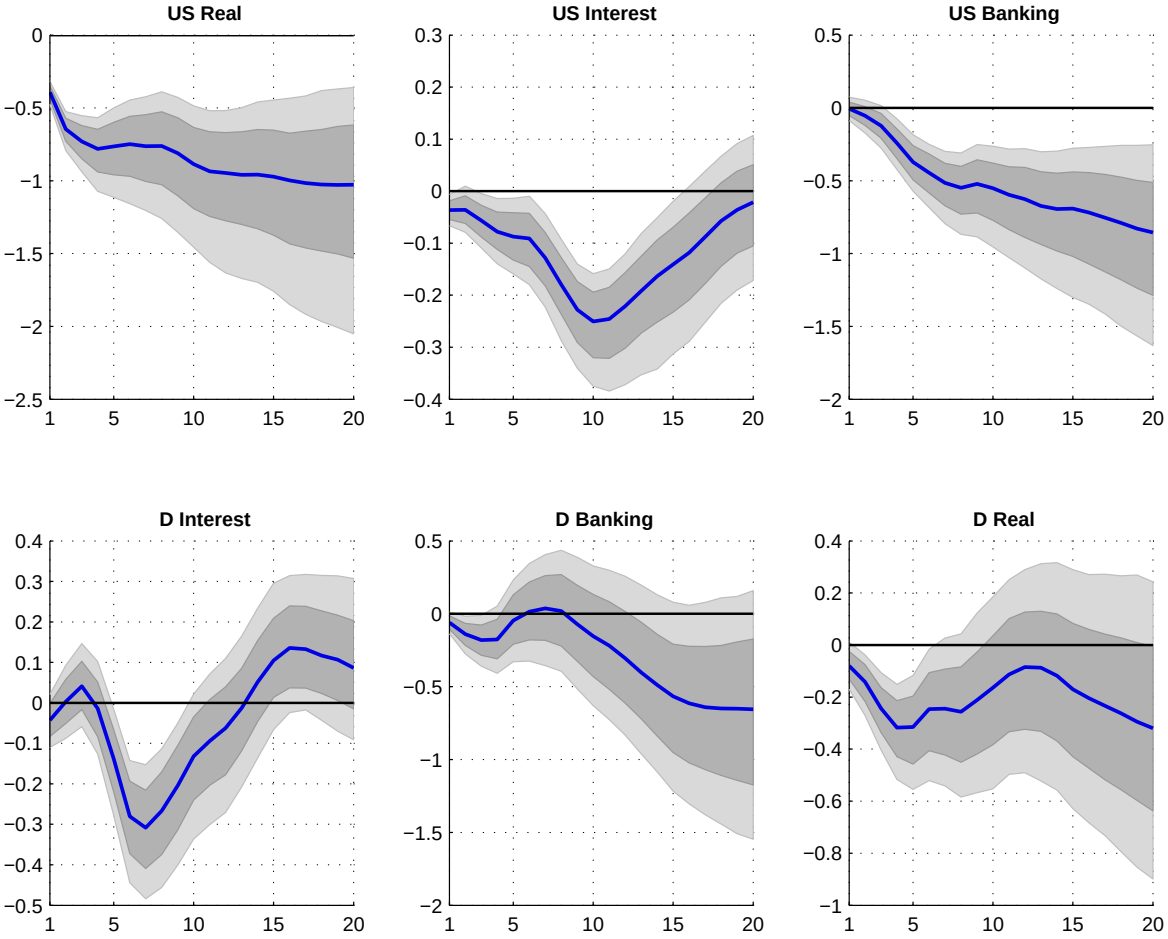


Figure 2: Responses of all variables to a contractionary shock of one standard deviation in size in the common component of U.S. real variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

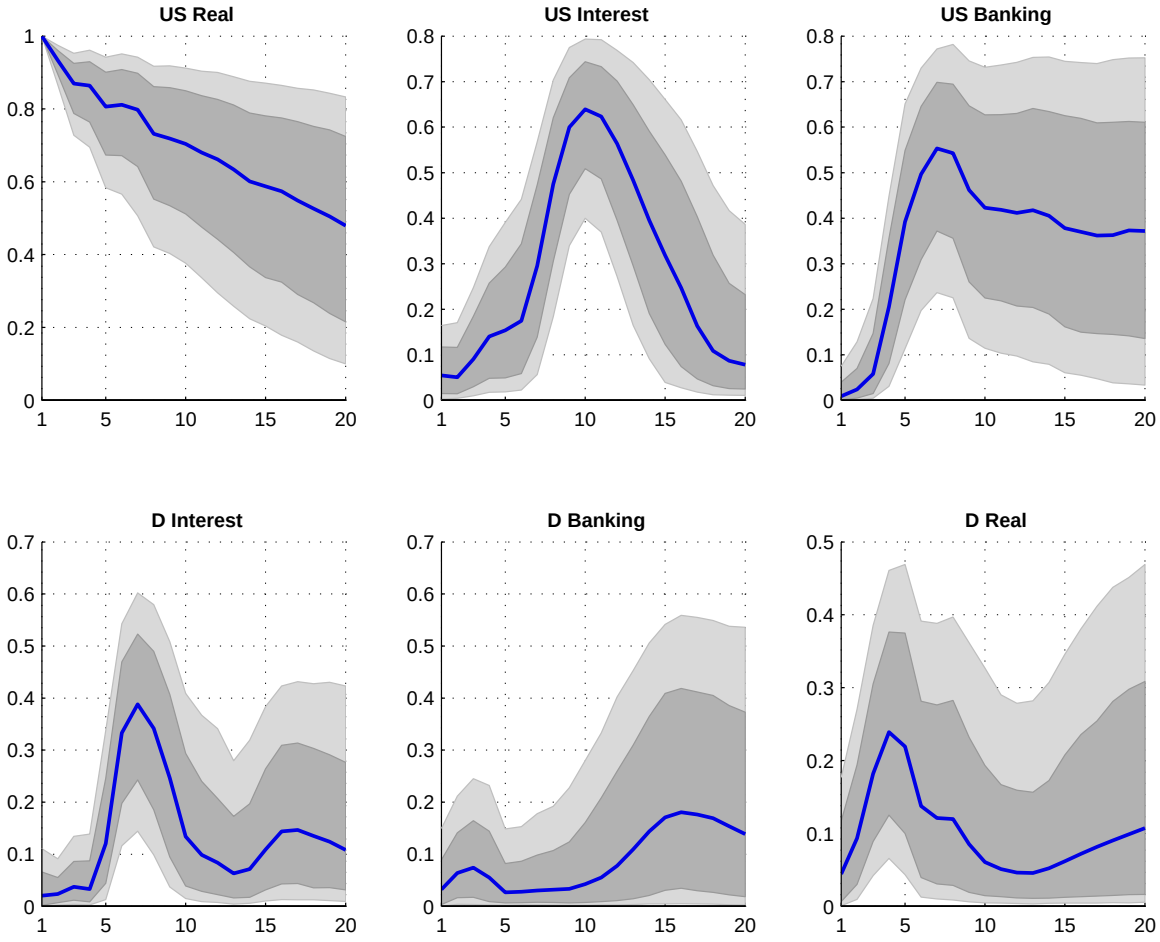


Figure 3: Fraction of the variance explained by a contractionary shock in the common component of U.S. real variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

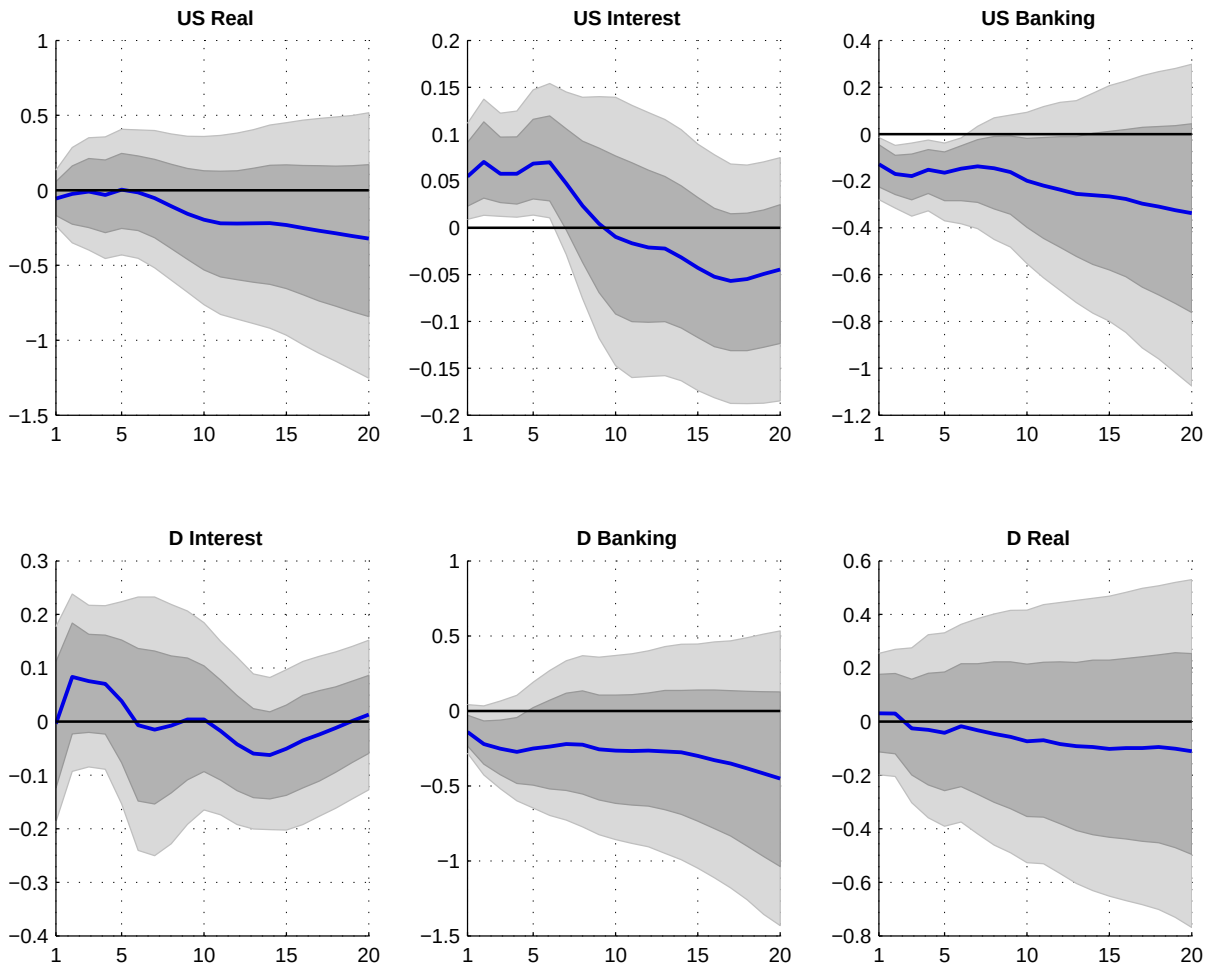


Figure 4: Responses of all variables to a contractionary nominal shock. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass. A sign restriction operates on the responses of the U.S. interest and banking factors for the first six months after the shock.

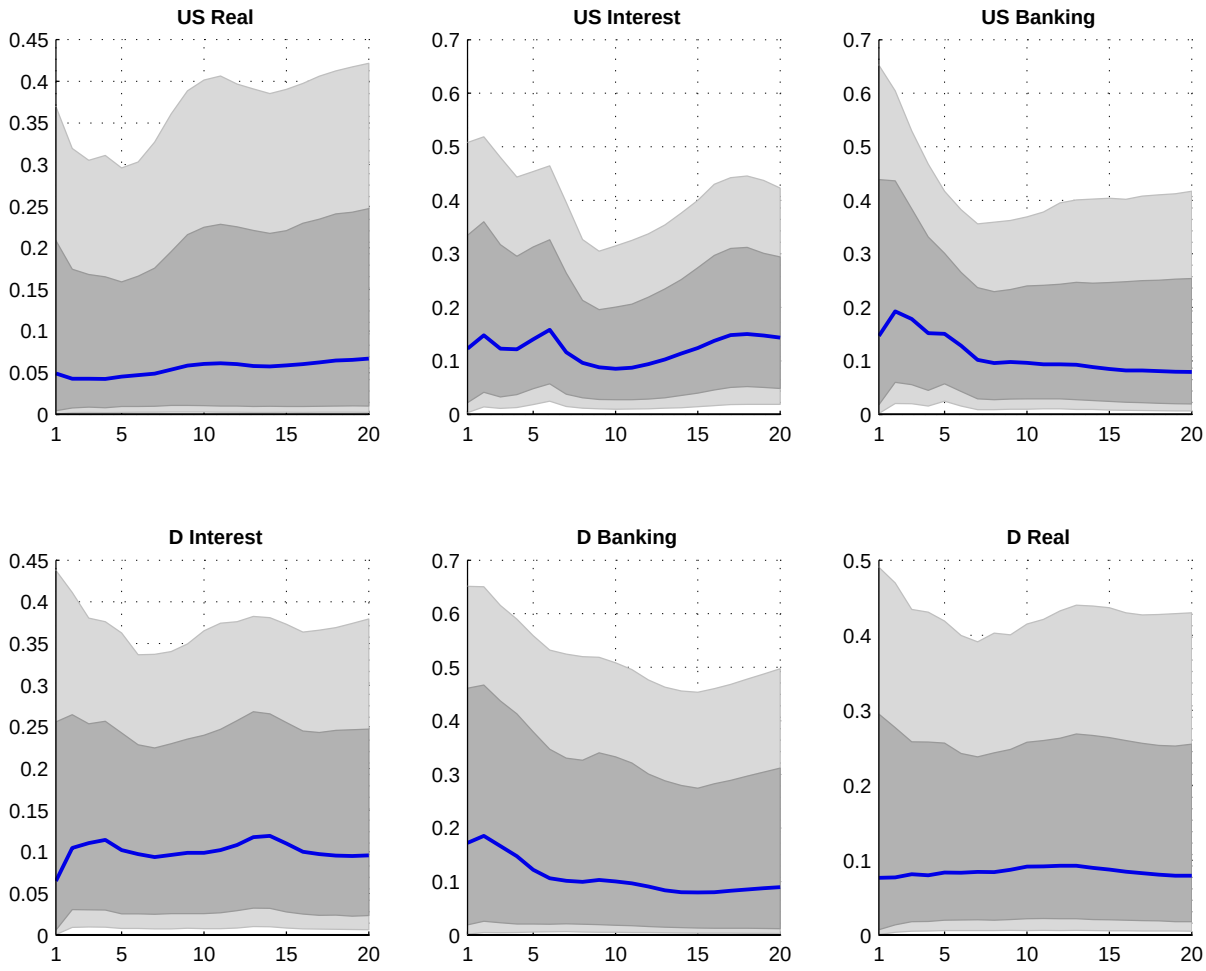


Figure 5: Fraction of the variance explained by a contractionary nominal shock. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

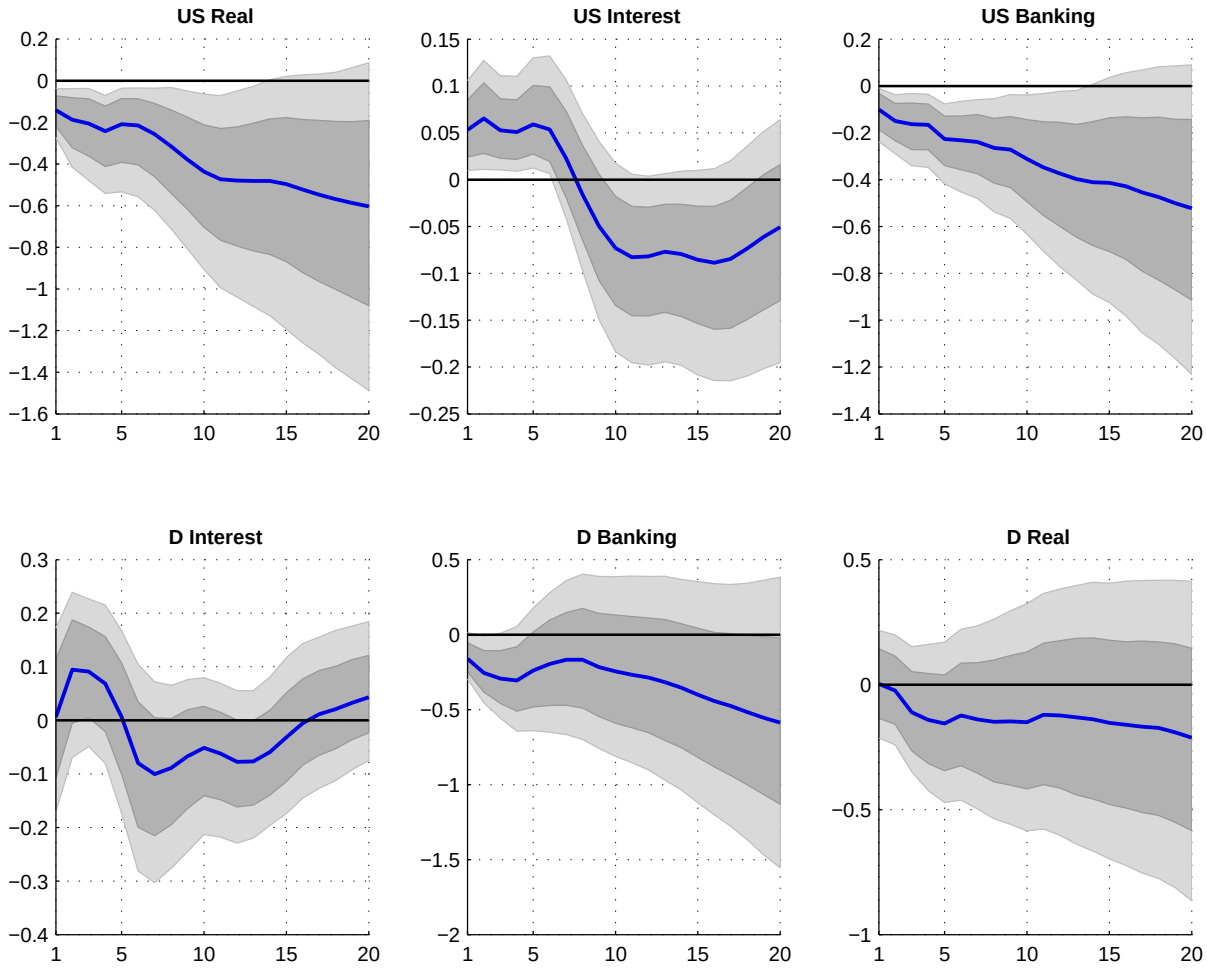


Figure 6: Responses of all variables to a contractionary shock of one standard deviation in size in the U.S. monetary factor. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass. A sign restriction operates on the responses of the U.S. real and the U.S. banking factors for the first six months after the shock.

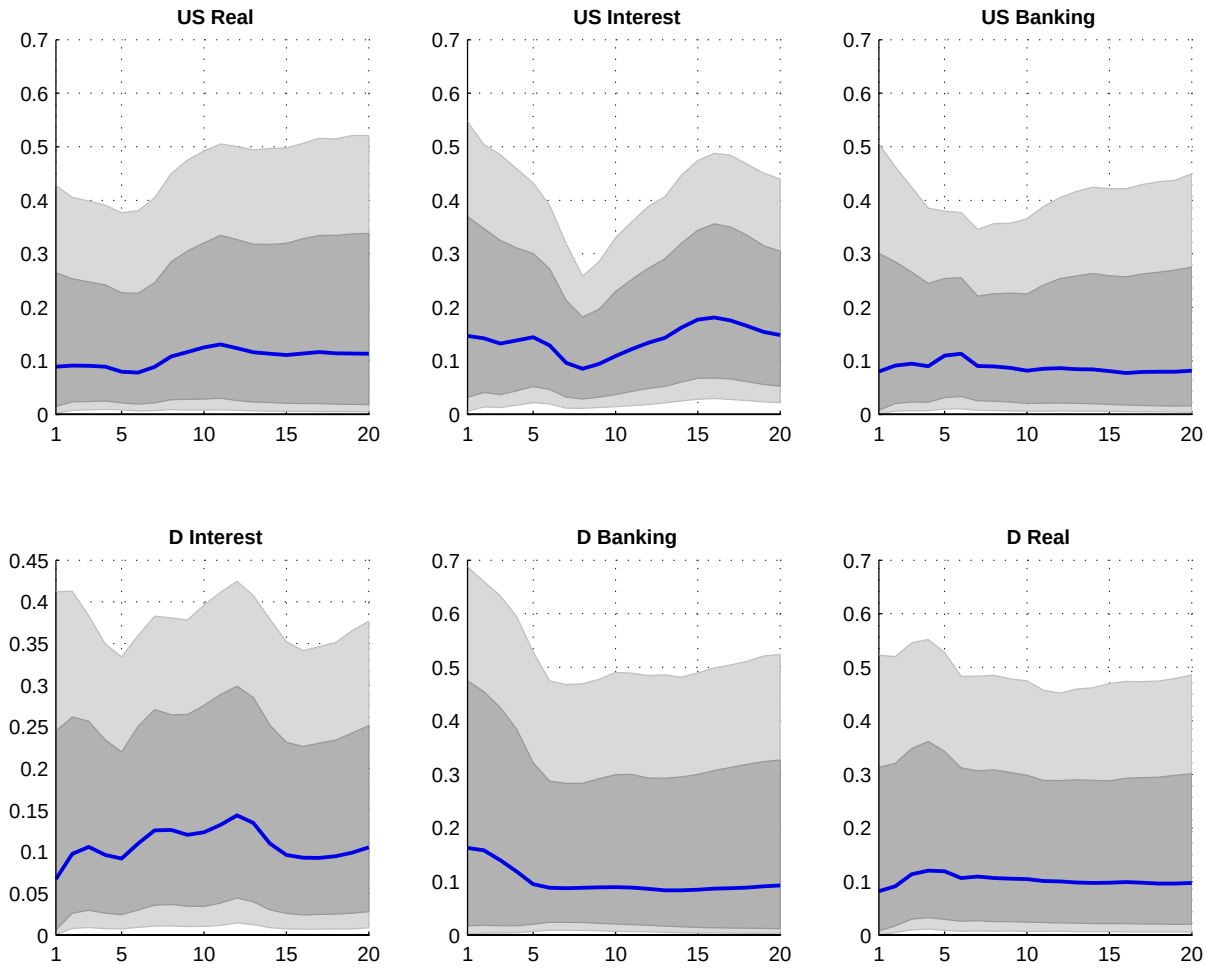


Figure 7: Fraction of the variance explained by a contractionary shock in the common component of U.S. interest rates. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

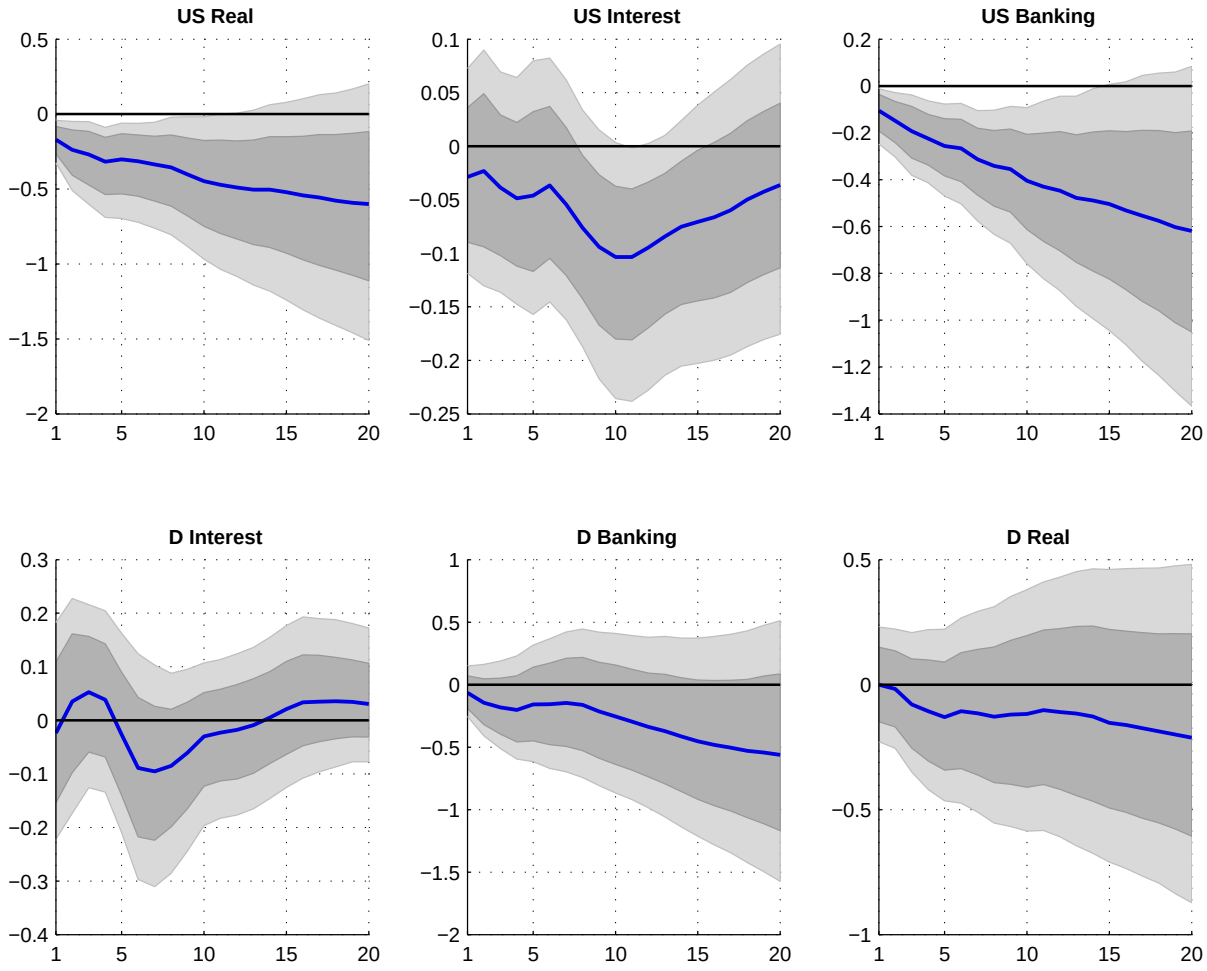


Figure 8: Responses of all variables to a contractionary shock of one standard deviation in size in the common component of U.S. financial variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass. A sign restriction operates on the responses of the U.S. real and the U.S. monetary factors for the first six months after the shock.

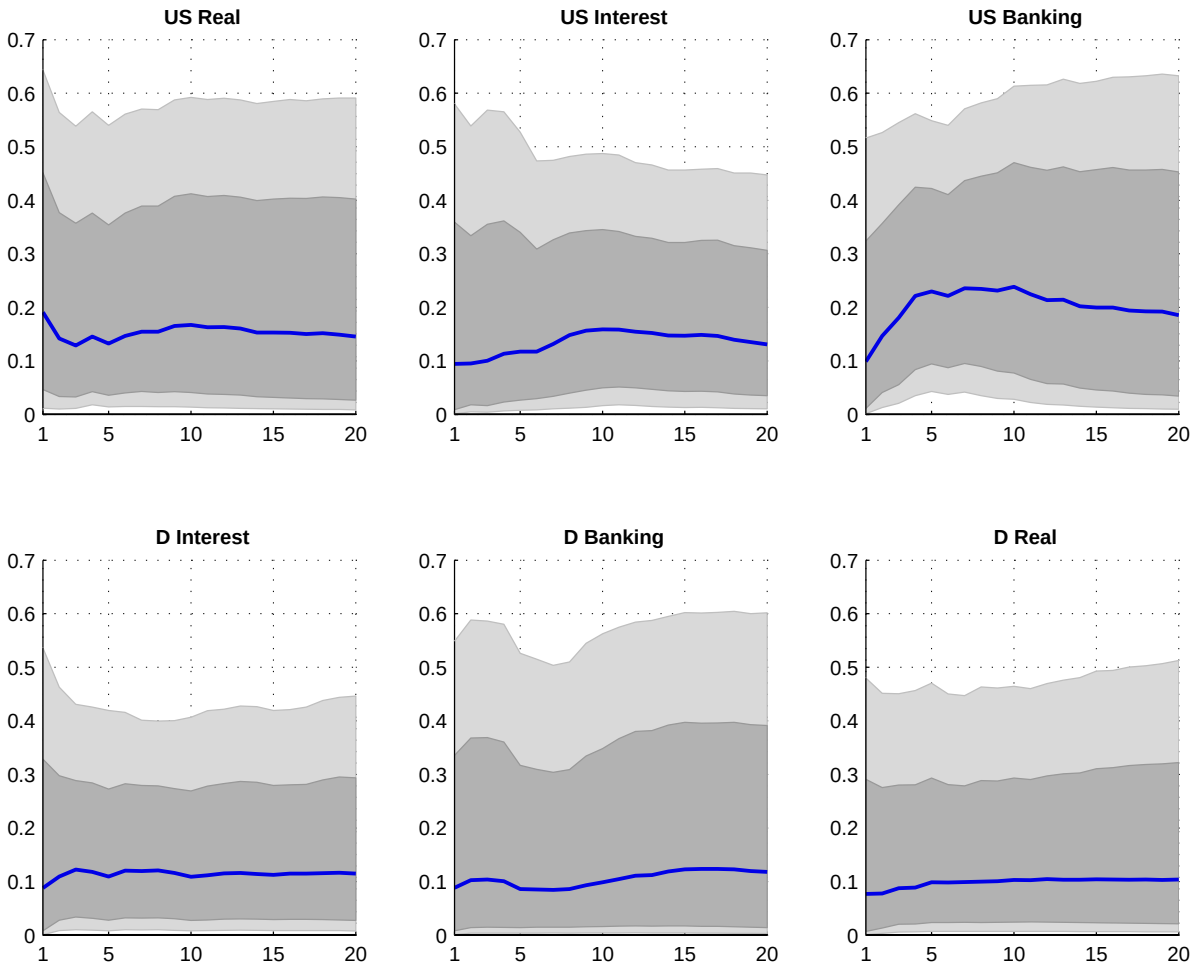


Figure 9: Fraction of the variance explained by a contractionary shock in the common component of U.S. financial variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

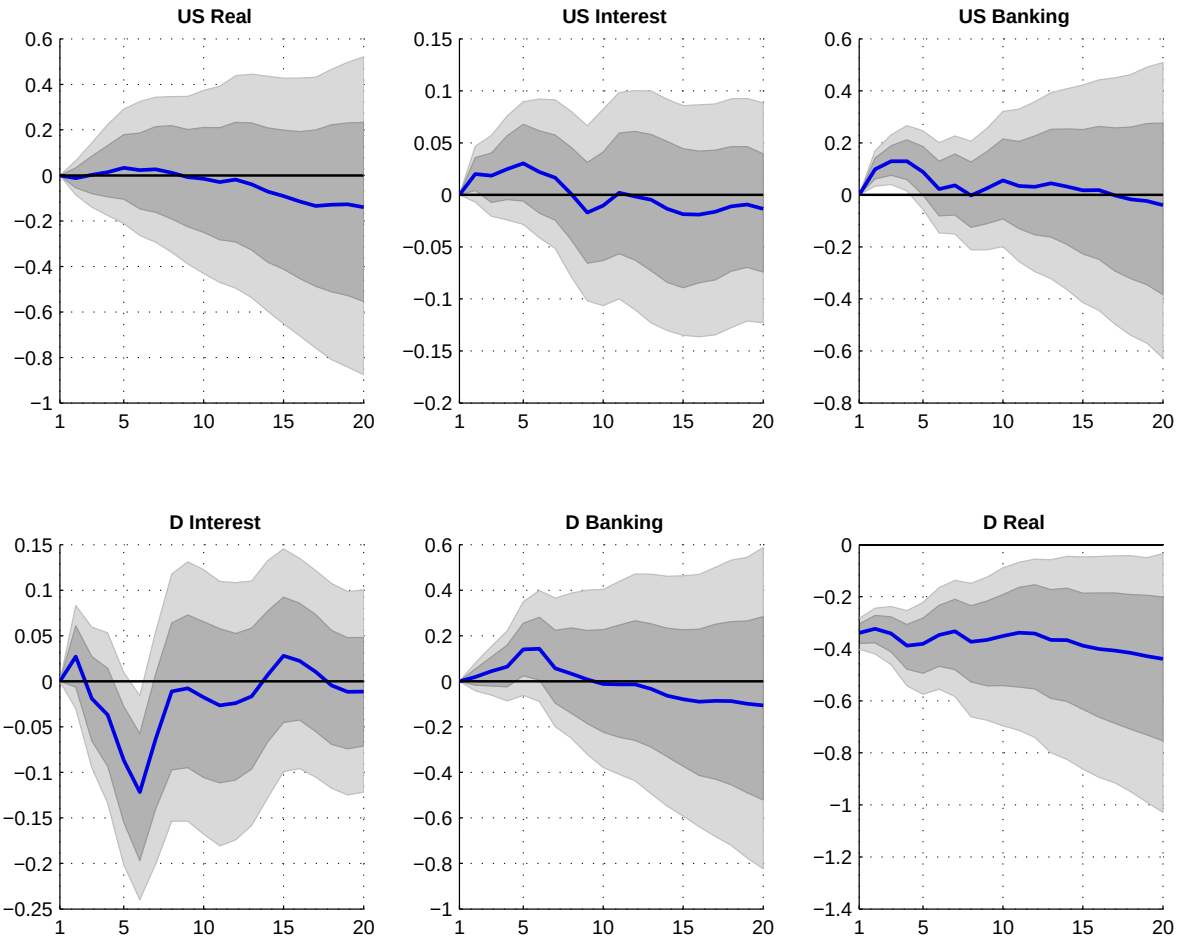


Figure 10: Responses of all variables to a contractionary shock of one standard deviation in size in the common component of German real variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

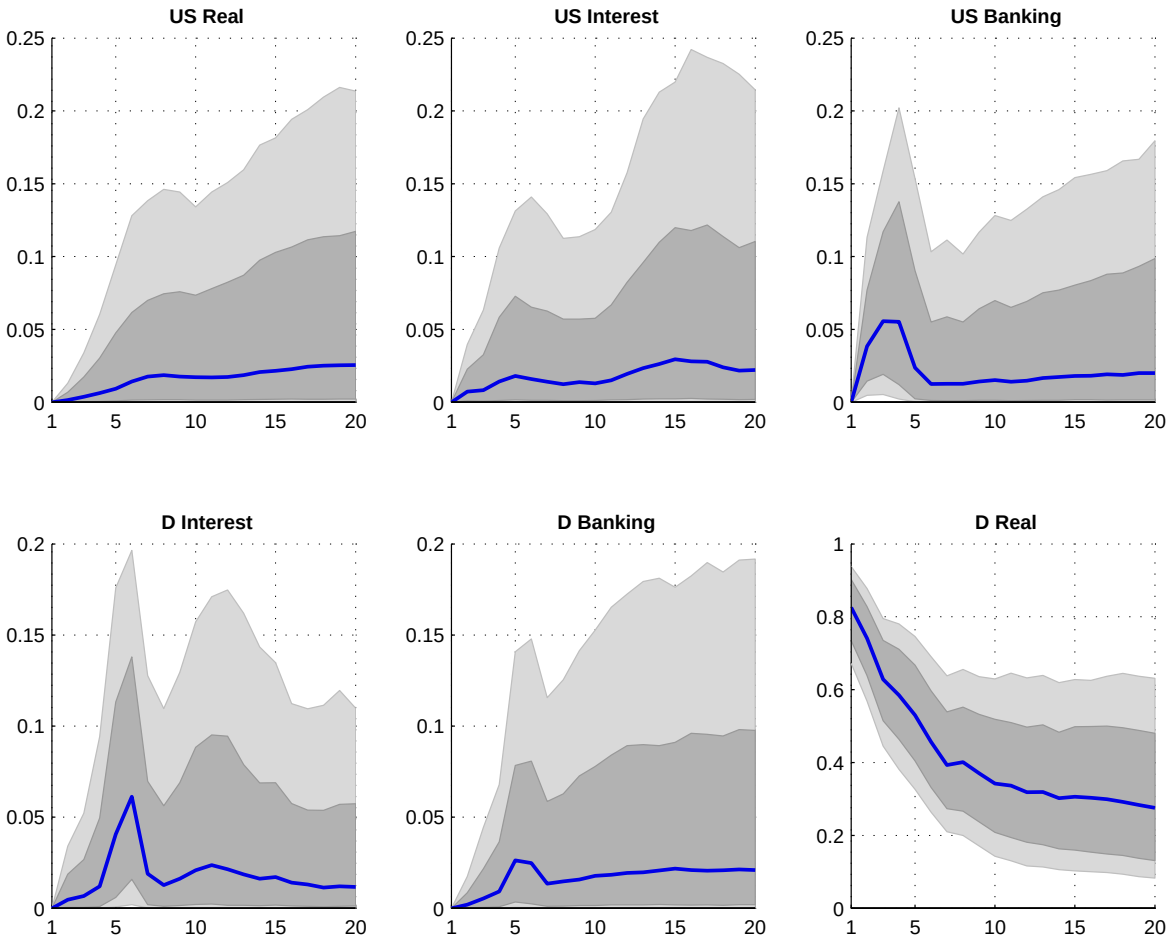


Figure 11: Fraction of the variance explained by a contractionary shock in the common component of German real variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

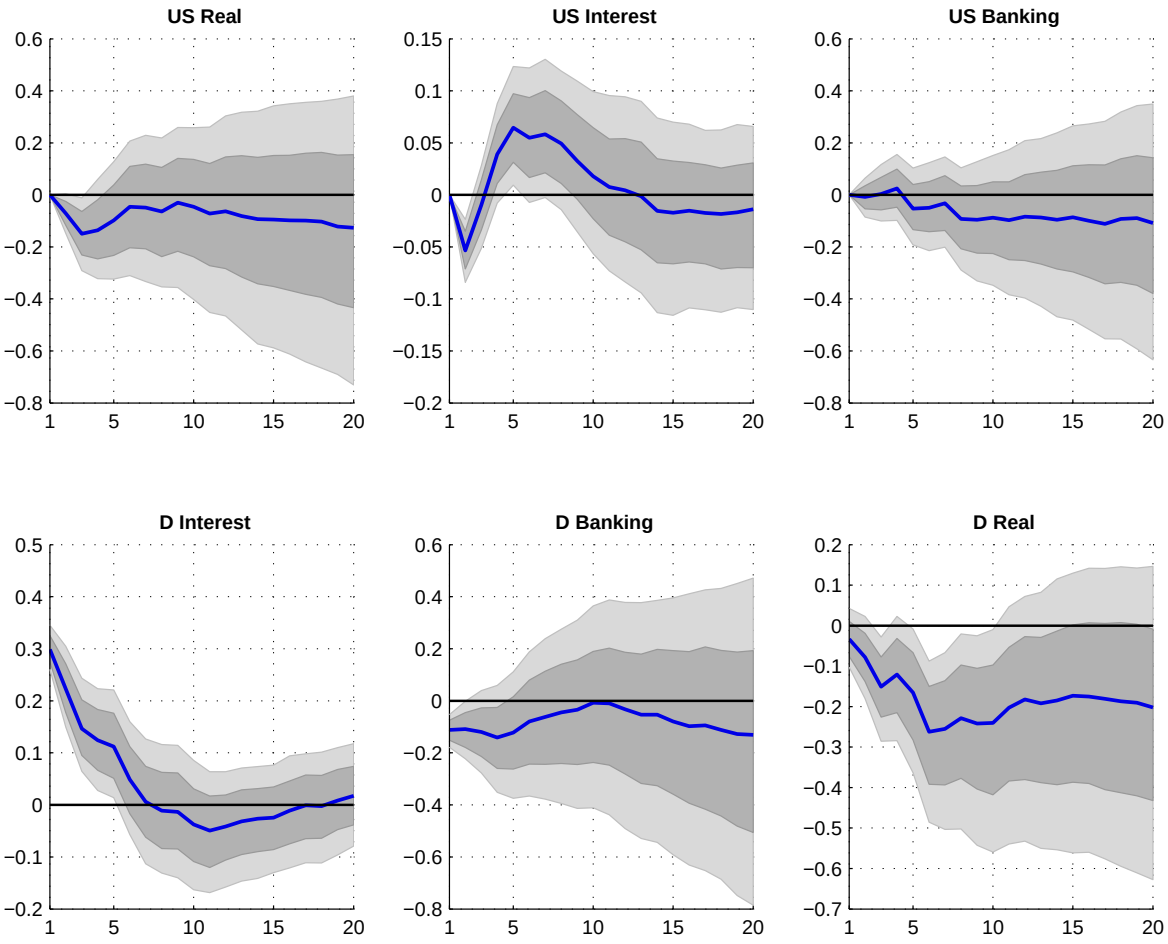


Figure 12: Responses of all variables to a contractionary shock of one standard deviation in size in the common component of German monetary variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

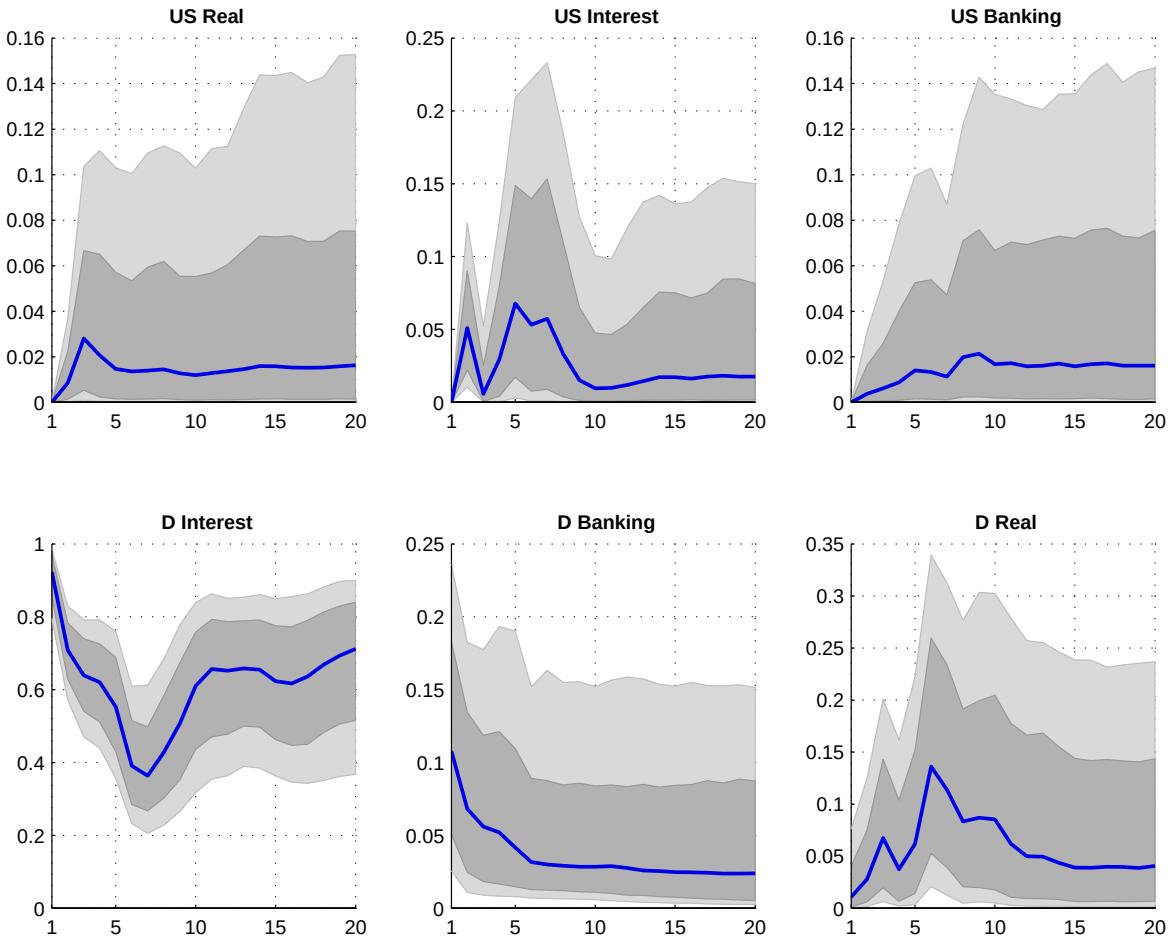


Figure 13: Fraction of the variance explained by a contractionary shock in the common component of German interest rates. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

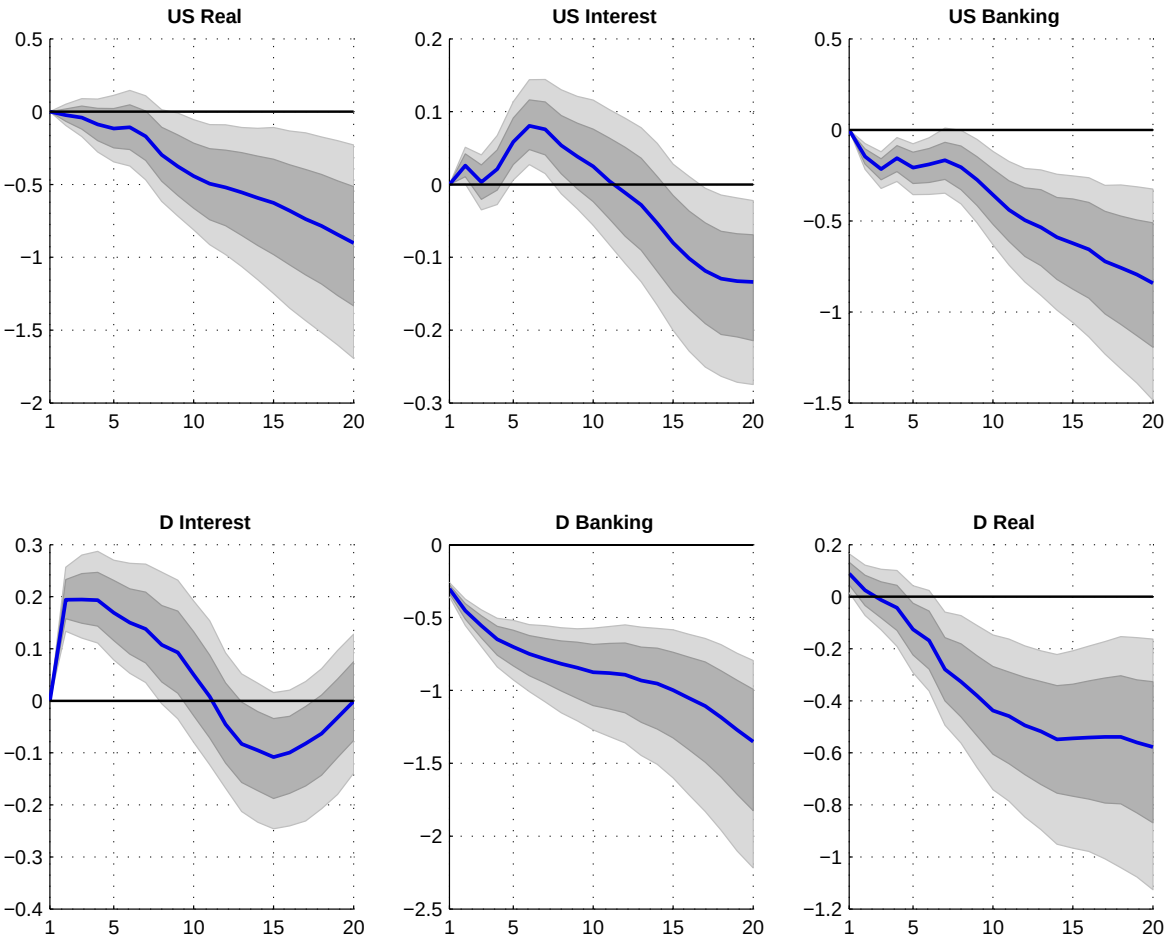


Figure 14: Responses of all variables to a contractionary shock of one standard deviation in size in the common component of German financial variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

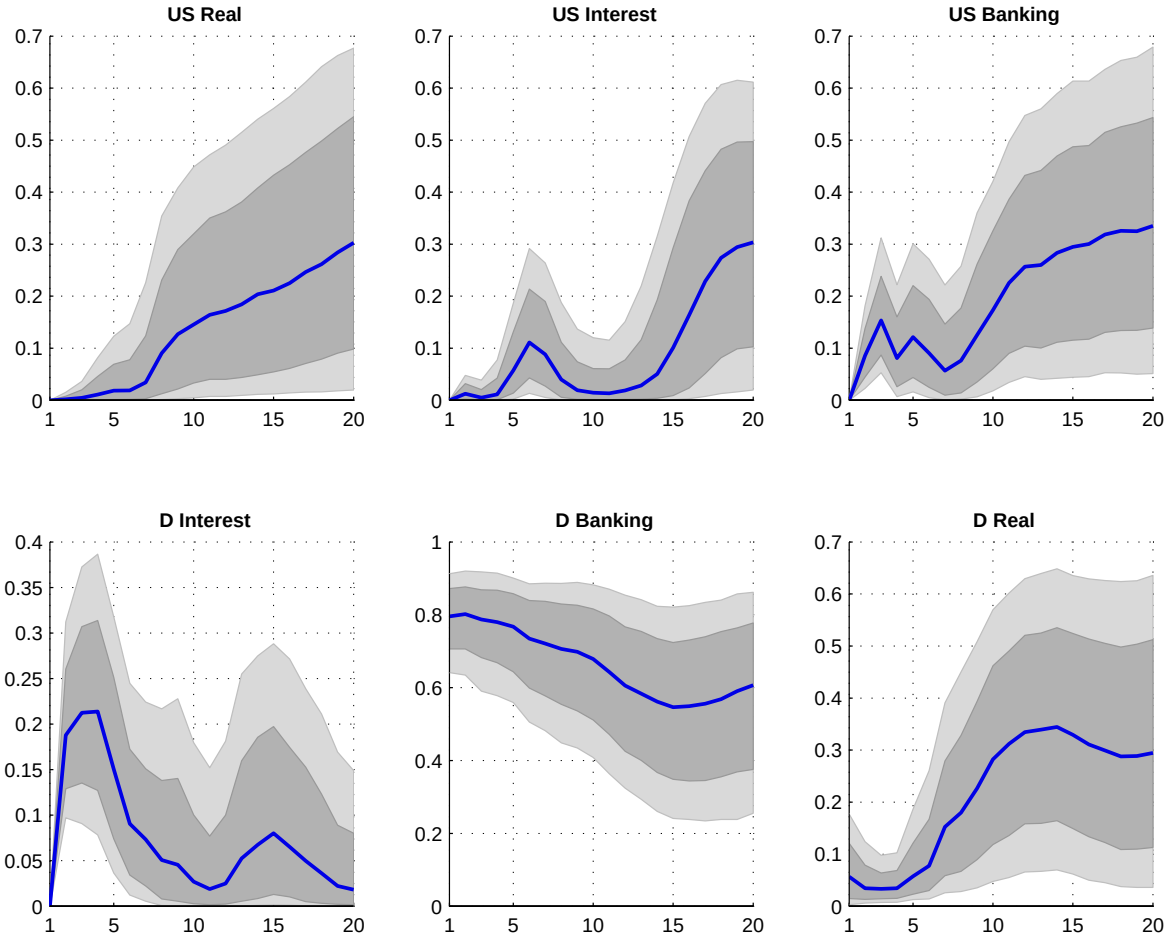


Figure 15: Fraction of the variance explained by a contractionary shock in the common component of German financial variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

C.3 Impulse Response Analysis (1925:9 to 1931:5)

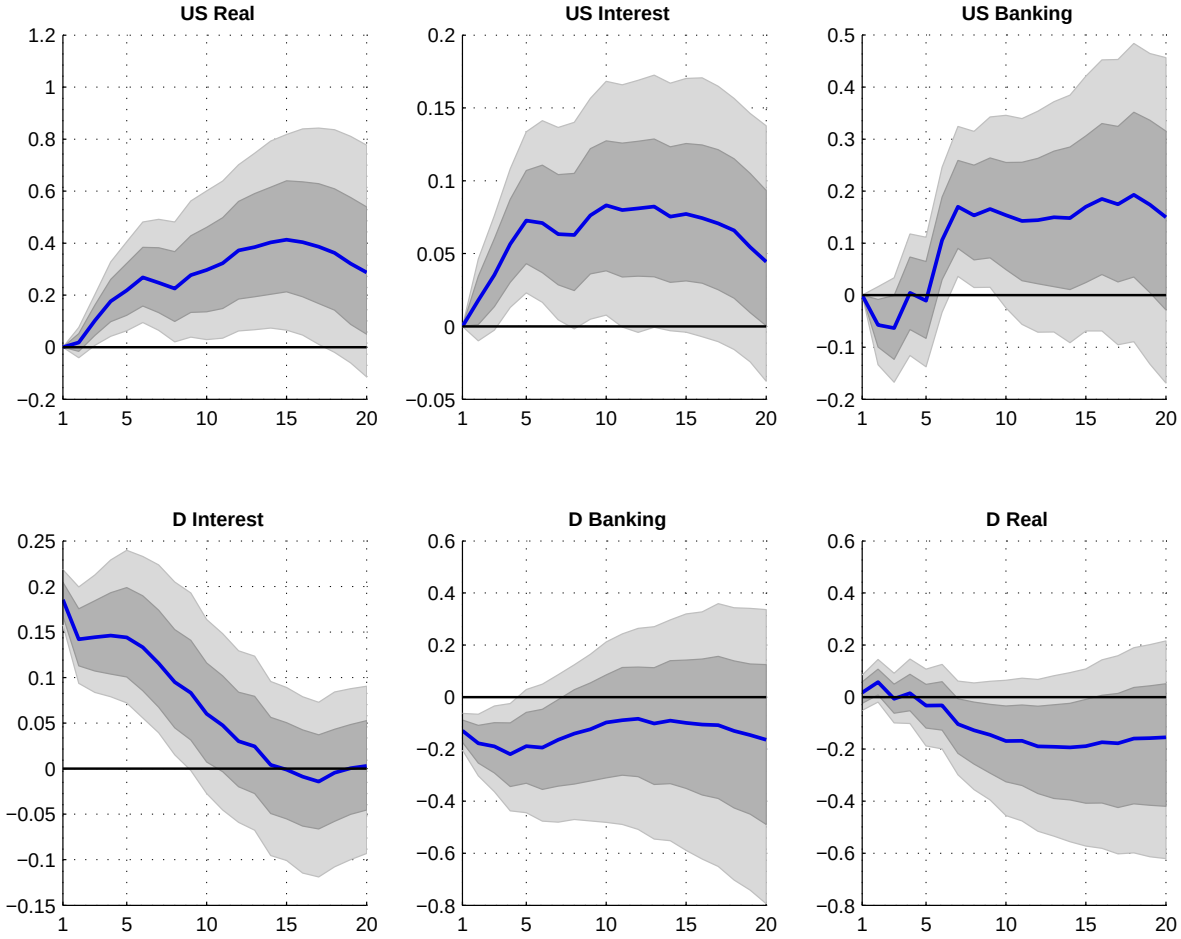


Figure 16: Responses of all variables to a contractionary shock of one standard deviation in size in the common component of German monetary variables when sample period is truncated to 1931:5. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

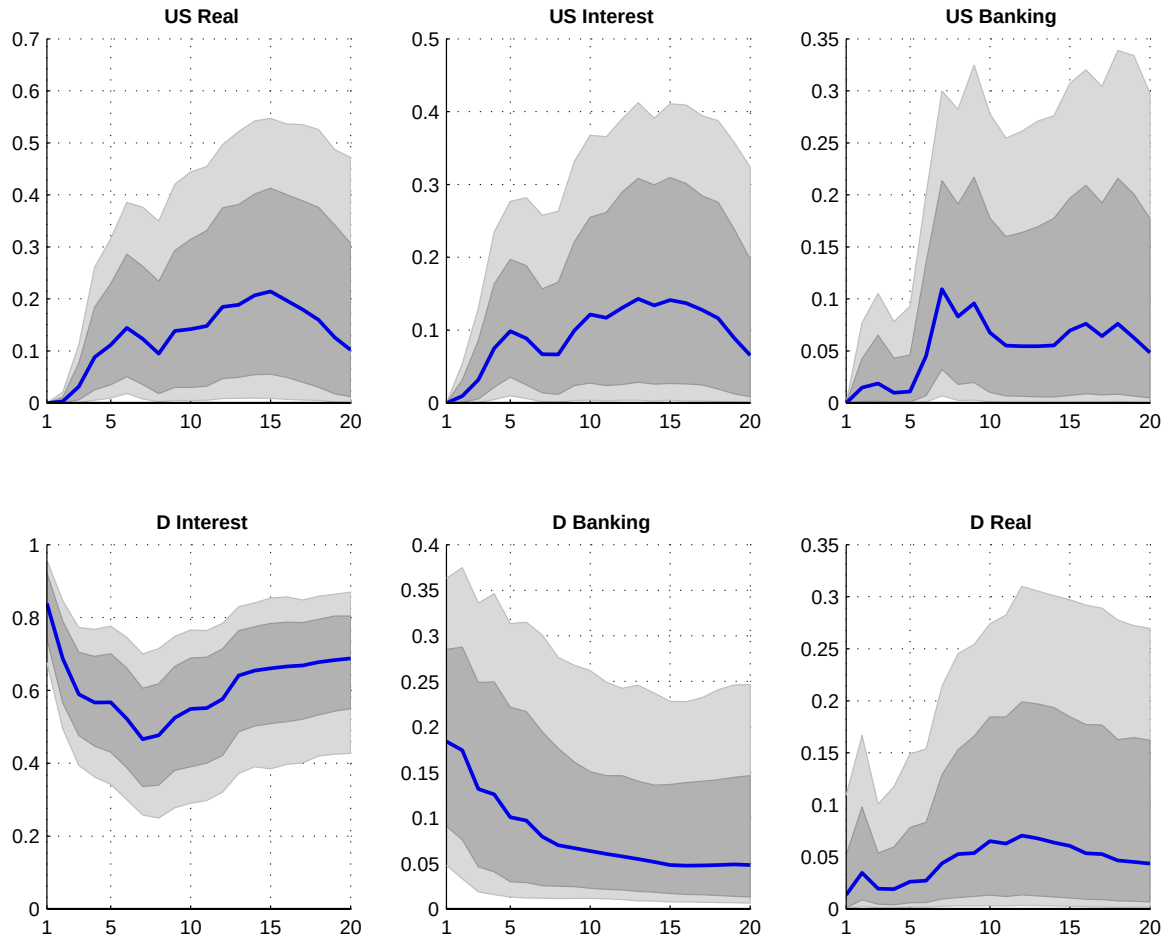


Figure 17: Fraction of the variance explained by a contractionary shock in the common component of German monetary variables when sample period is truncated to 1931:5. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

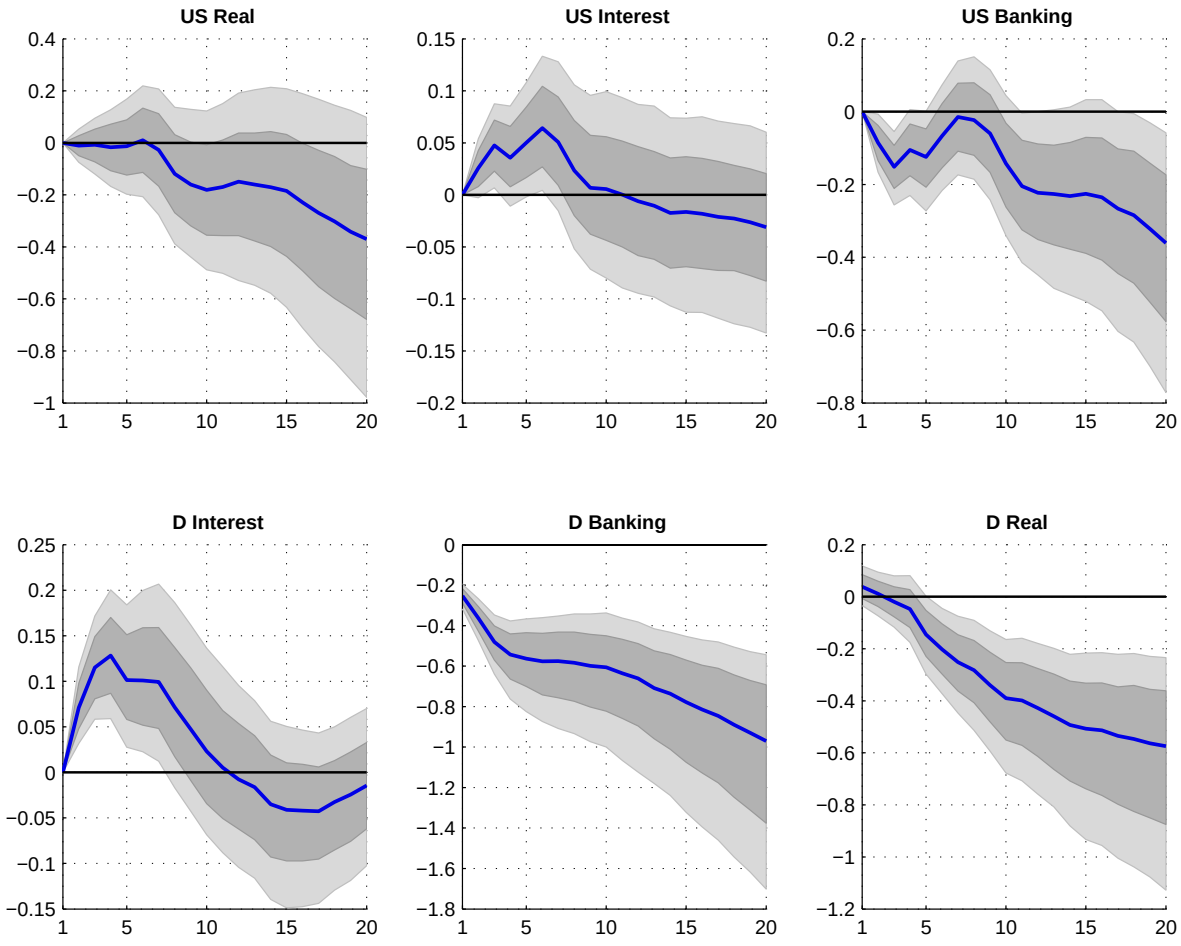


Figure 18: Responses of all variables to a contractionary shock of one standard deviation in size in the common component of German financial variables when sample period is truncated to 1931:5. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

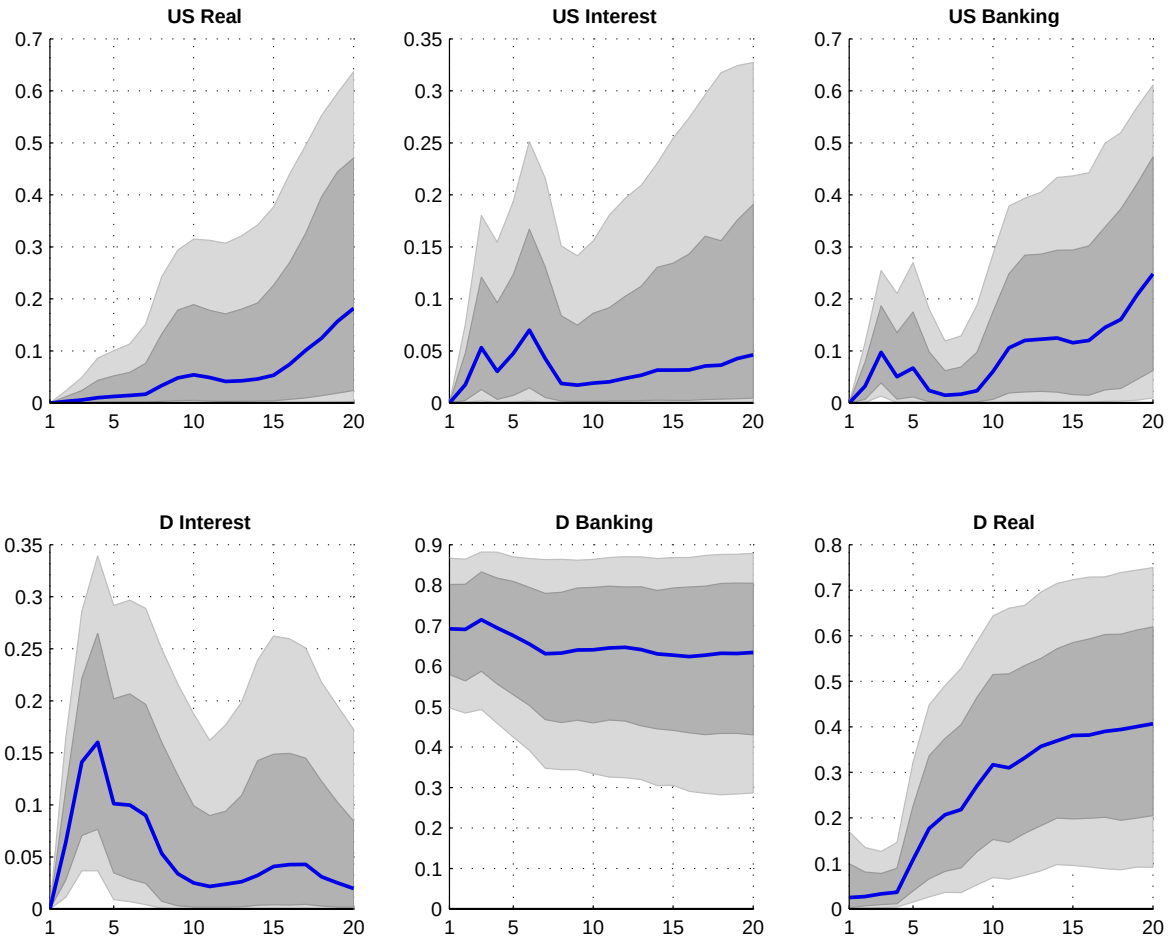
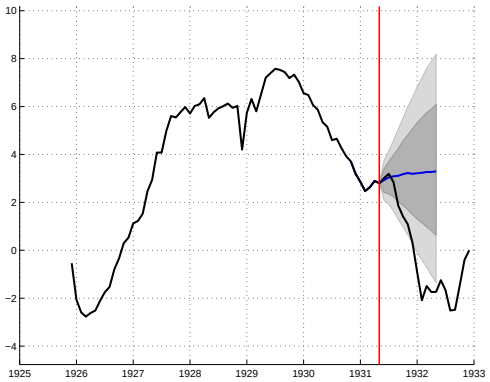


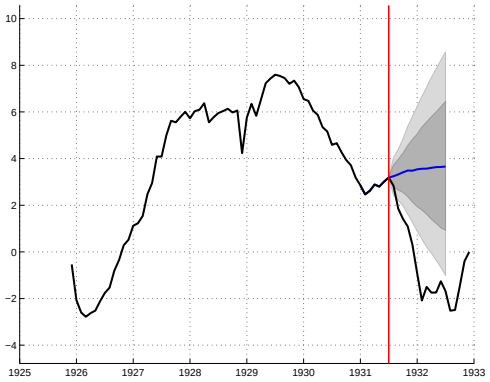
Figure 19: Fraction of the variance explained by a contractionary shock in the common component of German financial variables when sample period is truncated to 1931:5. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

C.4 Forecasting the Depression

(a) March 1931



(b) May 1931



(c) July 1931

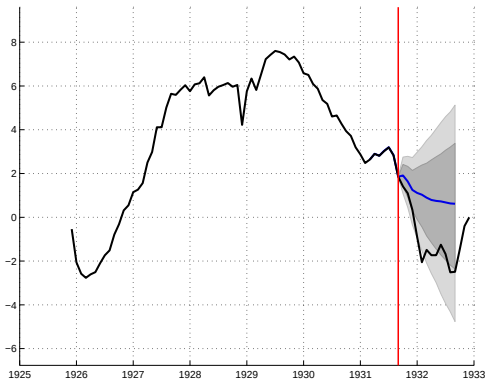
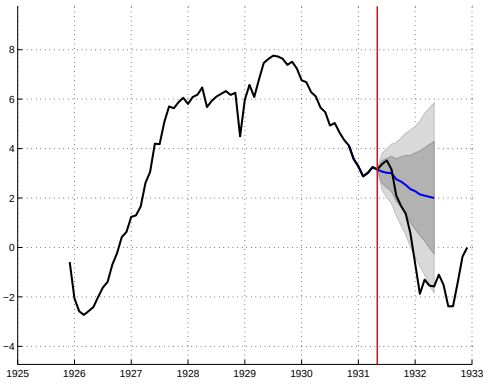
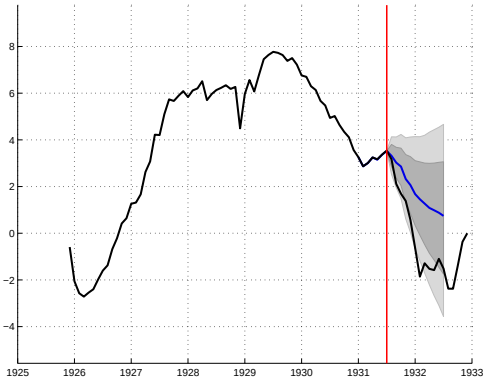


Figure 20: Forecasting the German real sector from March 1931, May 1931 and July 1931, using German real variables only. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

(a) March 1931



(b) May 1931



(c) July 1931

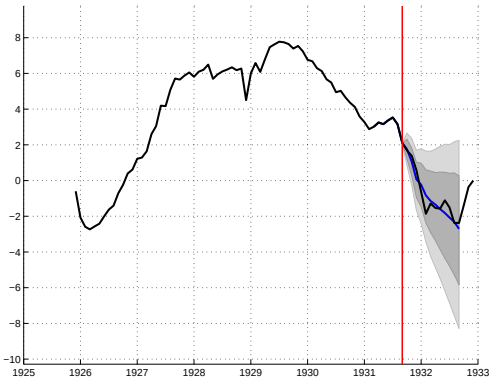
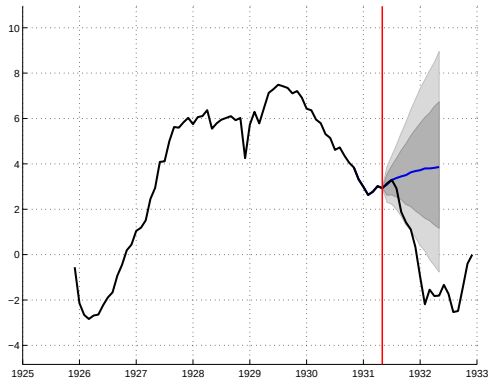
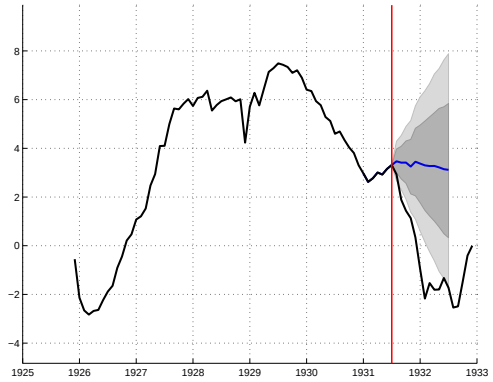


Figure 21: Forecasting the German real sector from March 1931, May 1931 and July 1931, using German real and banking variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

(a) March 1931



(b) May 1931



(c) July 1931

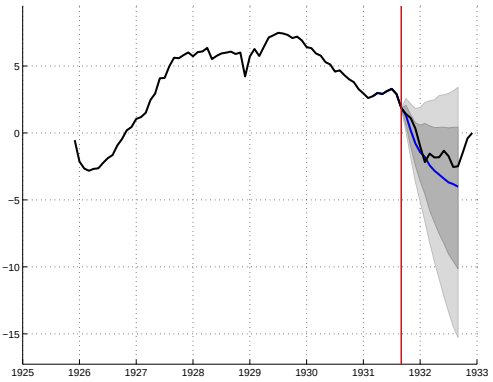
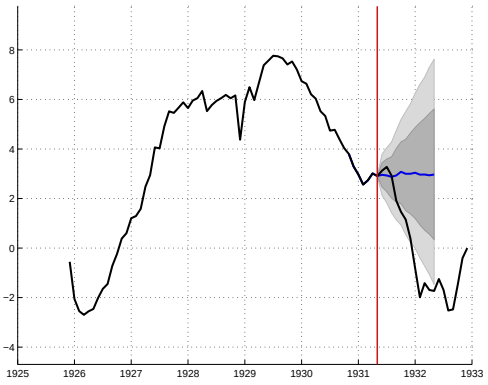
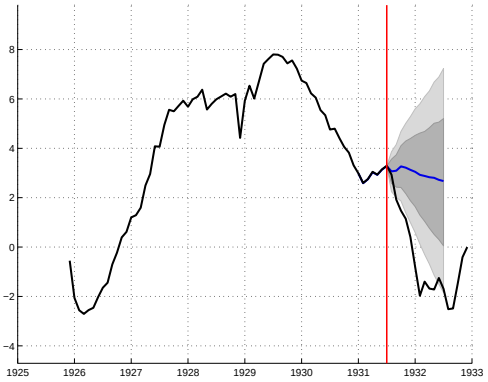


Figure 22: Forecasting the German real sector from March 1931, May 1931 and July 1931, using German real variables and interest rates. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

(a) March 1931



(b) May 1931



(c) July 1931

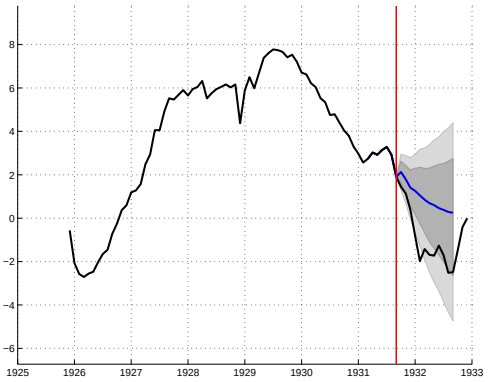
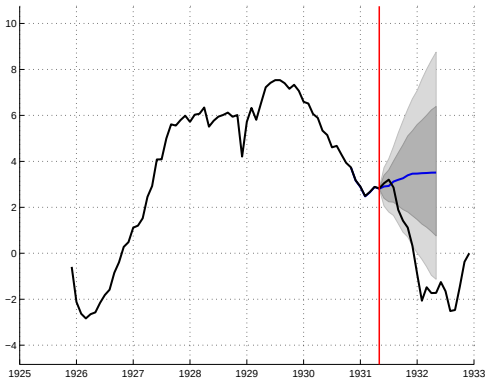
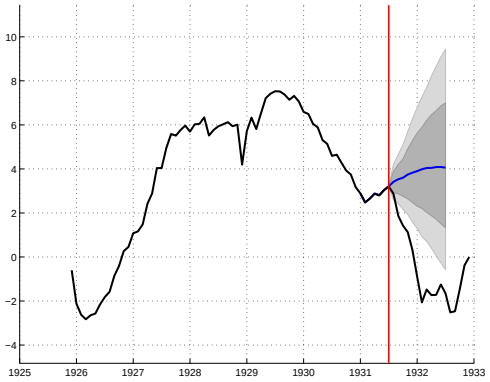


Figure 23: Forecasting the German real sector from March 1931, May 1931 and July 1931, using German and U.S. real variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

(a) March 1931



(b) May 1931



(c) July 1931

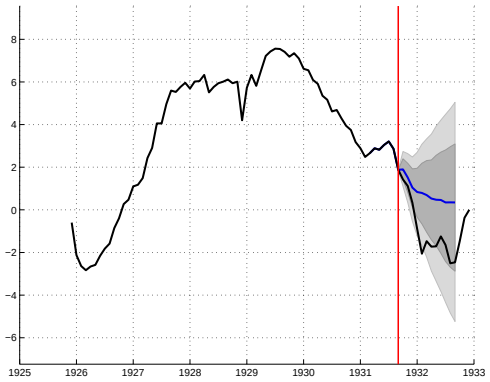
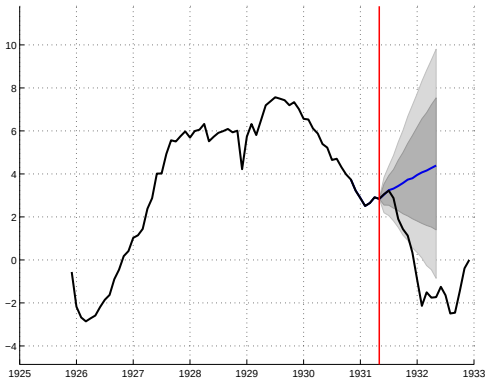
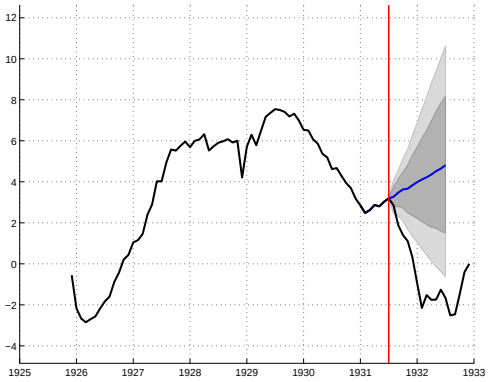


Figure 24: Forecasting the German real sector from March 1931, May 1931 and July 1931, using German real and U.S. banking variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

(a) March 1931



(b) May 1931



(c) July 1931

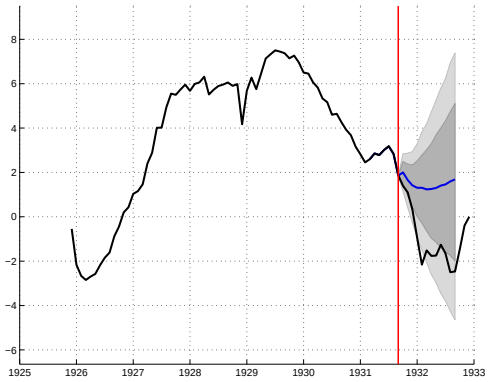
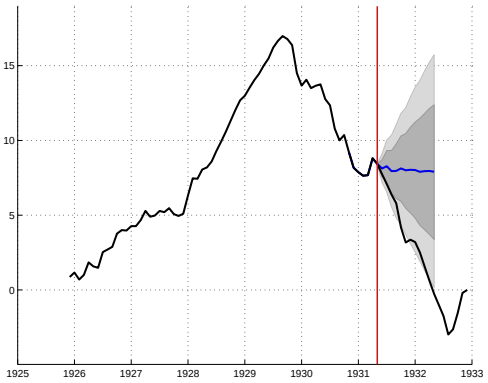
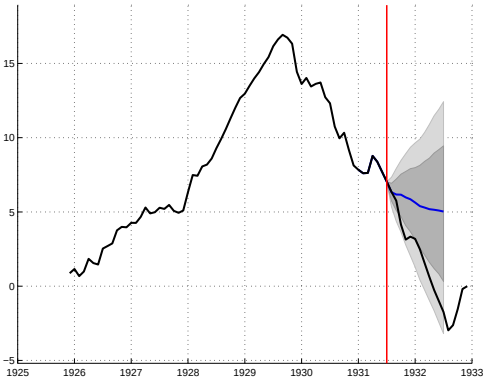


Figure 25: Forecasting the German real sector from March 1931, May 1931 and July 1931, using German real and U.S. interest rates. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

(a) March 1931



(b) May 1931



(c) July 1931

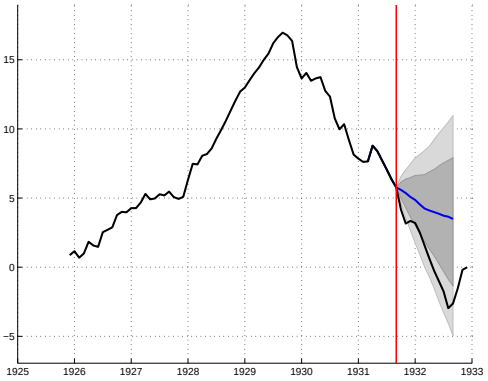
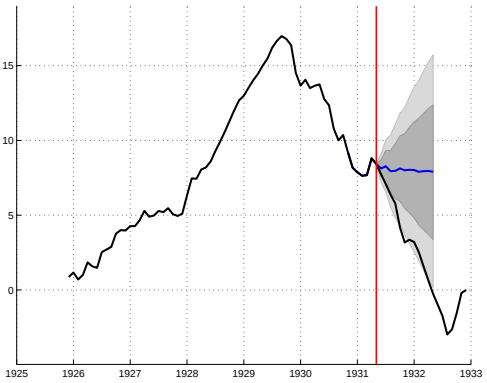
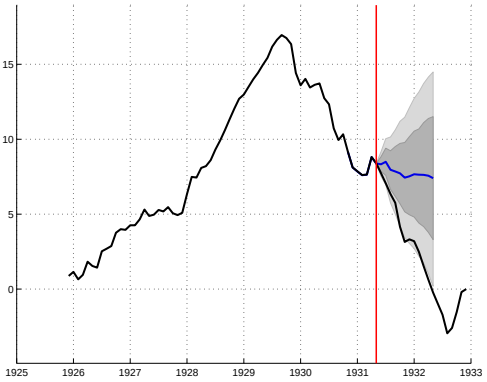


Figure 26: Forecasting the U.S. real sector from March 1931, May 1931 and July 1931, using U.S. real variables only. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

(a) Real Factor



(b) Real and Banking Factors



(c) Real and Monetary Factors

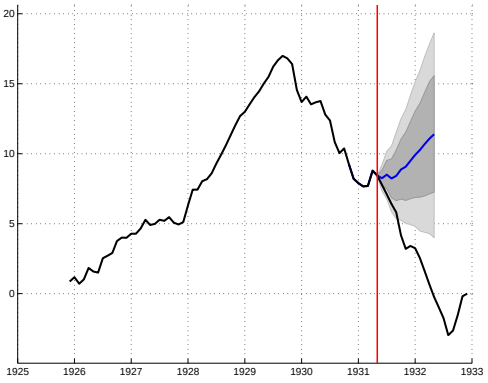
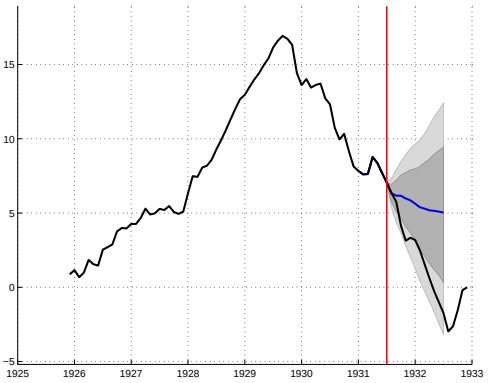
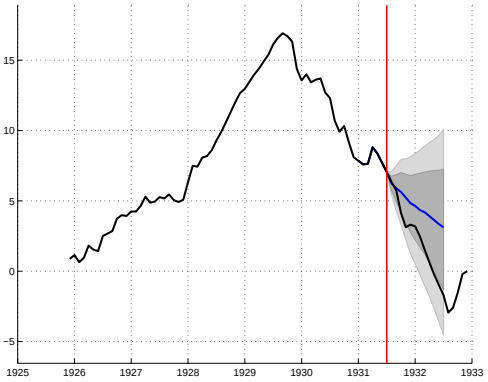


Figure 27: Forecasting the U.S. real sector from March 1931, using U.S. real, banking, and monetary variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

(a) Real Factor



(b) Real and Banking Factors



(c) Real and Monetary Factors

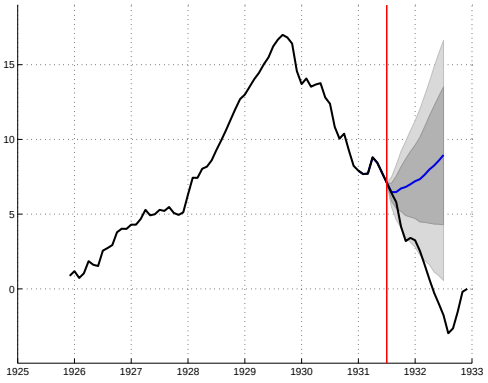
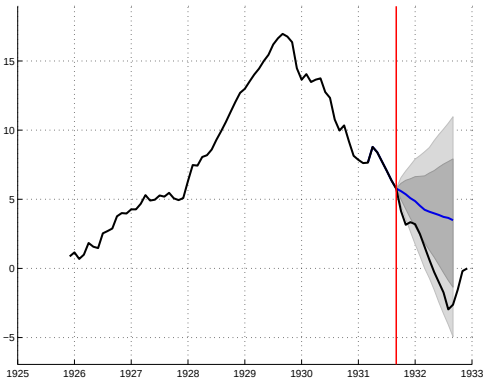
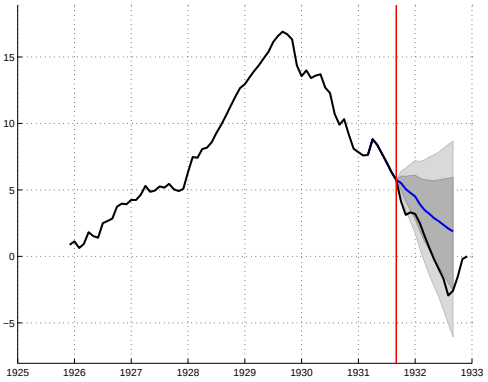


Figure 28: Forecasting the U.S. real sector from May 1931, using U.S. real, banking, and monetary variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

(a) Real Factor



(b) Real and Banking Factors



(c) Real and Monetary Factors

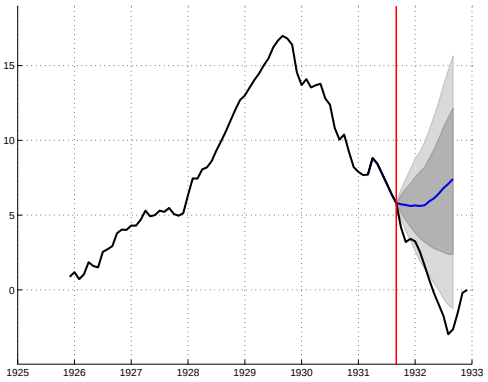
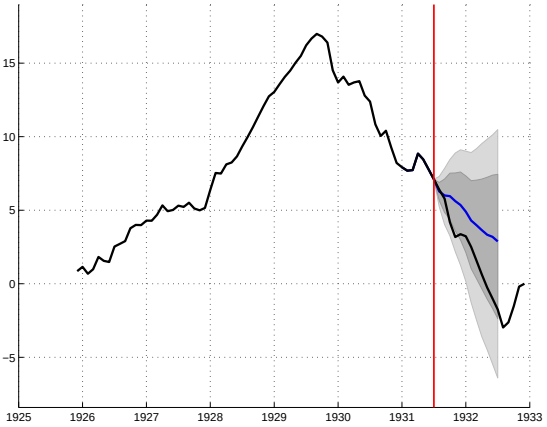


Figure 29: Forecasting the U.S. real sector from July 1931, using U.S. real, banking, and monetary variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

(a) U.S. Real and German Banking, May 1931



(b) U.S. Real and German Monetary, May 1931

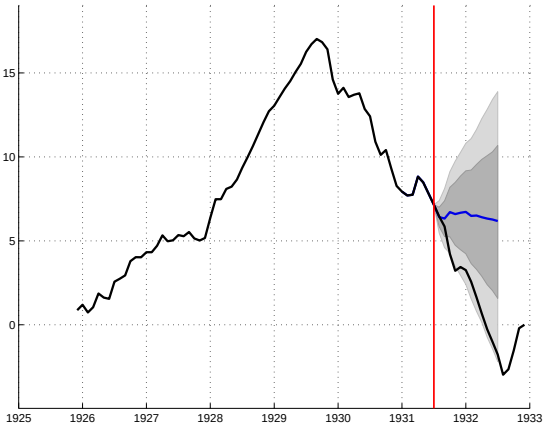
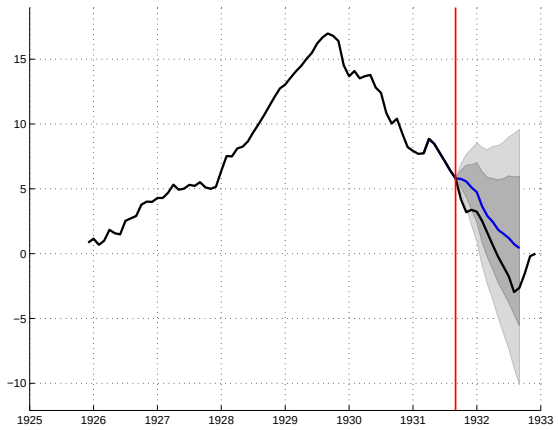


Figure 30: Forecasting the U.S. real sector from May 1931, using German banking and monetary variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

(a) U.S. Real and German Banking, July 1931



(b) U.S. Real and German Monetary, July 1931

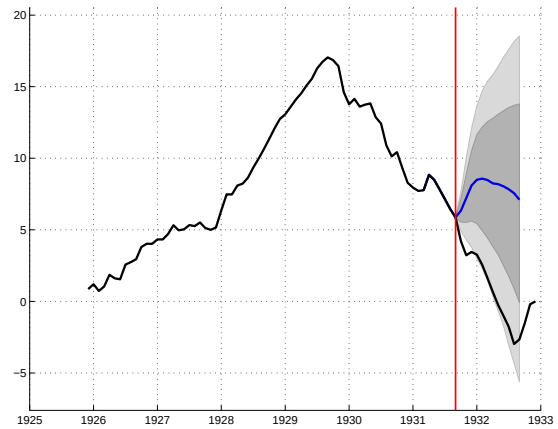


Figure 31: Forecasting the U.S. real sector from July 1931, using German banking and monetary variables. The dark gray shaded area represents 68% and the light shaded area 90% of the posterior probability mass.

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