On the low-frequency relationship between public deficits and inflation*

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Abstract

We estimate the low-frequency relationship between fiscal deficits and inflation and pay special attention to its potential time variation by estimating a time-varying VAR model for U.S. data from 1900 to 2011. We find the strongest relationship neither in times of crisis nor in times of high public deficits, but from the mid-1960s up to 1980. Employing a structural decomposition of the low-frequency relationship and further narrative evidence, we interpret our results such that the low-frequency relationship between fiscal deficits and inflation is strongly related to the conduct of monetary policy and its interaction with fiscal policy after World War II.

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1 Introduction

The recent economic crisis has led to an enormous increase in fiscal deficits. Concerns that high deficits are inflationary in the medium or long run originate in economic theory developed by Sargent and Wallace (1981) and Leeper (1991). These studies share the idea that not fiscal policy alone, but rather its interaction with monetary policy, is key for the inflationary consequences of public deficits. In particular, Leeper (1991) shows that, when outstanding government debt is not backed by future primary surpluses and at the same time the monetary authority does not apply the Taylor principle, eventually the price level increases to revalue the amount of real debt.[1] This theory is referred to as the fiscal theory of the price level (FTPL). What is more, Cochrane (2001) and Bianchi and Melosi (2013) show that the FTPL implies that the relationship between inflation and fiscal stance can be especially strong at lower frequencies. While Sargent and Wallace (1981) stresses a different channel of public deficits' inflationary consequences, e.g. by seignorage, their "unpleasant monetarist arithmetic" shares the importance of lower frequencies in the relationship between public deficits and inflation.

Employing the FTPL mechanism, Bianchi and Ilut (2012) explicitly account for time variation in the interaction between monetary and fiscal policy by estimating a structural Markov-switching dynamic stochastic general equilibrium (MS-DSGE) model. The authors consider time series from 1955 and show that the rise of inflation in the 1960s and 1970s as well as the fall that occurred in the early 1980s can be explained by a change in the interaction of fiscal and monetary policy. Furthermore, the authors find that this policy change also accounts for breaks in the persistence and volatility of inflation and breaks in the link between inflation and fiscal variables.

In this paper, we take the finding by Bianchi and Ilut (2012) as motivation to re-investigate the relationship between the fiscal stance and inflation taking into account that this relationship potentially varies over time. We employ a time-varying parameter Vector-Autoregression (TVP-VAR) model, which requires compared to the MS-DSGE model fewer parametric restrictions. The TVP-VAR allows us to consider long time series for the U.S. from 1900 to 2011. Moreover, since the studies by Cochrane (2001), Bianchi and Melosi (2013), and Sargent and Wallace (1981) suggest that the relationship could be most pronounced beyond business cycle frequencies, we investigate the relationship at lower frequencies.

The investigation of the relationship between the fiscal stance and inflation across the

[1] According to the Taylor principle, the monetary authority should aggressively fight inflation by raising the nominal interest rate by more than the increase in inflation above target, which leads to an increase in the real interest rate.

long time span is beneficial in many ways. Firstly, it includes periods of even higher debt growth than the one the U.S. faces today. Secondly, it contains periods such as the Great Depression, which are of current interest as they also involved financial crises. Thirdly, it comprises episodes of different policy regimes which are accompanied by different degrees of central bank independence or different kinds of fiscal-monetary policy regimes. We choose the framework of a TVP-VAR model because it accommodates the different time episodes and structural changes in a single empirical framework. The setup of the TVP-VAR model and the computation of the low-frequency relationship closely follows Sargent and Surico (2011). We measure fiscal stance by real debt growth minus the gross real interest rate, which is defined as primary deficits over one-period lagged debt (see Sims, 2011). For the sake of readability, we refer to primary deficits over one-period lagged debt as deficits over debt throughout the paper.

We find a positive low-frequency relationship between public deficits over debt and inflation for the U.S. which is time-varying. In the first half of the 20th century, the low-frequency relationship is volatile and only partly positive. After WWII, the relationship is stable and significantly positive up to 1980. It reaches its highest value in the years between 1973, the end of the Bretton Woods system, and 1979, the year in which Paul Volcker became Chairman of the Federal Reserve System. After 1980, the low-frequency relationship is stable and insignificantly different from zero.

The findings regarding the time after 1973 are in line with conventional wisdom. The period of the 1970s is usually characterized either by a central bank not responding strongly to inflation (e.g. Clarida, Gali, and Gertler, 2000; Lubik and Schorfheide, 2004) or by a central bank which has lost its ability to control inflation (Sims, 2011), while the fiscal authority was playing a dominant role (e.g. Davig and Leeper, 2007; Bianchi and Ilut, 2012; Bianchi and Melosi, 2013). Meltzer (2010, p. 485) states that the Federal Reserve had "accepted its role as a junior partner by agreeing to coordinate actions with the administration's fiscal policy" up to 1979. As Meltzer (2010) points out, the conduct of policy changes with the chairmanship of Paul Volcker in the fourth quarter of 1979. Volcker rebuilt much of the independence and credibility which the Federal Reserve had lost over the two proceeding decades (see, e.g., Taylor, 2011).

In order to interpret the reduced form results structurally, we determine how much of the movement in the low-frequency relationship is due to monetary policy shocks. The monetary policy shock is identified by a recursive identification scheme as in Sims (1992) and Christiano, Eichenbaum, and Evans (1996). We find that the increase in the low-frequency relationship in the 1970s is due to the transmission mechanism of monetary policy shocks. Using a counterfactual analysis, we show that the breakdown of the low-frequency relationship at the beginning of 1980 is related to a structural change in the economy regarding the transmission mechanism of monetary policy shocks. Conversely, if this switch had not occurred, the low-frequency relationship between public deficits and inflation would have been high for the last 30 years, too. Our finding is line with the theories on the interaction of monetary and fiscal policy. These theories demonstrate that the transmission mechanism of monetary policy shocks depends on the interaction of monetary and fiscal policy. Consequently, we interpret our findings as confirmation that the low-frequency relationship between fiscal deficits and inflation is strongly related to the independence of monetary policy and its interaction with fiscal policy. In a policy regime where the central bank accommodates the action of the fiscal authority, and the fiscal authority is willing to accept high inflation rates, high deficits are associated with inflation in the long run.

The relationship between fiscal deficits and inflation has been studied extensively. Studies employing constant coefficient models which do not focus on the low frequency of the time series find no relationship between fiscal deficits and inflation or seignorage respectively for the U.S. after WW II. (Niskanen, 1978; McMillin and Beard, 1982; King and Plosser, 1985; Joines, 1985; Giannitsarou and Scott, 2008) However, another strand of the literature with a more international focus finds a significant relationship for high-inflation countries (see, for example Catão and Terrones, 2005; Lin and Chu, 2013) in the long run. To the best of our knowledge, the present paper is the first which estimates the time-varying low-frequency relationship by investigating long time series for the United States employing a TVP-VAR model.

The remainder of the paper is organized as follows. Section Two illustrates how we measure the low-frequency relationship in a simple two-variable approach for different subsamples. In Section Three, we present the TVP-VAR framework and its estimation. Section Four presents and discusses the estimation results regarding the low-frequency relationship: In this section we also isolate the conduct of monetary policy as the factor driving the low-frequency relationship from 1960 onward. Section Five concludes.

2 The measurement of the low-frequency relationship

In this section, we introduce our two main variables, inflation and primary deficits over one-period lagged debt, and describe how we measure the low-frequency relationship between them. For the sake of readability, we denote the latter variable deficits over debt instead of primary deficits over one-period lagged debt throughout the paper.

Both time series range from 1875 to 2011. Following Sargent and Surico (2011), inflation (π) is measured as year-to-year first differences of the logarithmic GDP deflator, where the data is taken from the FRED II database starting in 1947Q1 and from Balke and Gordon (1986) before then. As a measure of the fiscal stance, we consider the variable deficits over debt (d). This measures debt growth minus the gross real interest rate. In contrast to the debt over output ratio or debt growth, this measure is not influenced by variables which are not controlled directly by the fiscal authority, such as output or the real interest rate. In order to gain intuition for the measure of fiscal stance, consider the opposite of our measure of fiscal stance – government's primary surplus over one-period lagged debt. This summarizes the net payments to bondholders either through interest rates or through retirement of bonds. Thus, a change in the deficits over debt measures the change in the fiscal authority's future liabilities. Furthermore, choosing this variable allows us to interpret our measure of the low-frequency relationship. If the measure is equal to one, the increase in primary deficit is matched by an increase in inflation and, thus, the real debt remains constant.

The time series for primary deficit and government debt held by the public are taken from Bohn (2008). The time series is of annual frequency. Since the remaining time series employed in this paper are of quarterly frequency, we decide to interpolate the annual data using the cubic-spline approach. Additionally, the time series for government debt is only available in par values and not in market values. Since we are interested in the low-frequency relationship, temporary differences between market and par values should not be critical (see also Bohn, 1991). However, we discuss the robustness of our results regarding interpolation and market value of debt in addition to other changes of specifications and assumptions in more detail in Section 4.2.

We employ the same approach as Lucas (1980) and Sargent and Surico (2011) to gauge the low-frequency relationship between two variables. In particular, this measure can be interpreted as the slope in a scatter plot of the low-frequency components of two filtered time series. To illustrate the measure, we filter inflation and deficits over debt by the filter suggested by Lucas (1980). Denote the unfiltered time series by x and the corresponding filtered time series by $x(\beta)$. The filter is defined as $x(\beta)_t = \alpha \sum_{k=-n}^{n} \beta^{|k|} x_{t+k}$, where $\alpha = (1-\beta)^2 / (1-\beta^2 - 2\beta^{n+1}(1-\beta))$ is chosen such that the sum of weights equals one. We set n to eight and choose $\beta = 0.95$, which achieves the intention of focusing on low-frequency variations. In Figure 1 we plot the filtered time series. The plot suggests that there are periods in which both time series share a co-movement. Figure 2(a) shows a scatter plot for the two time series. To keep the plot clearly arranged, we consider only the first quarter of each year. Using different quarters or even all observations does not affect the illustrative results.

To investigate potential time variation in the data, we divide the sample into four sub-

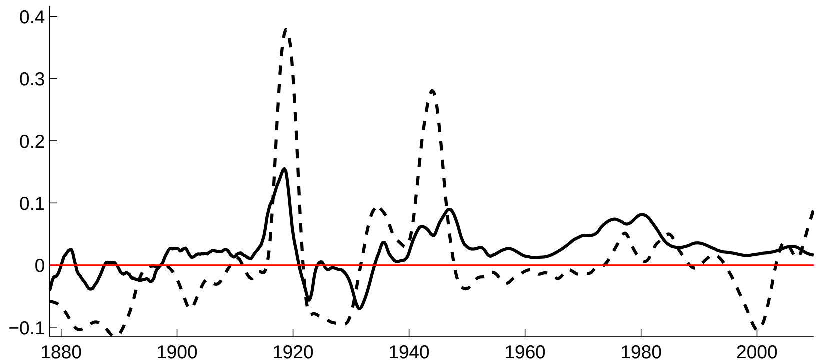


Figure 1: Filtered time series of inflation (solid) and deficit over debt (dashed). $\beta = 0.95$

samples: 1900-1933, 1934-1951, 1952-1983, and 1984-2009. The distinctive events for the sub-samples are the New Deal policy in 1934 and the abandonment of the Gold Standard one year earlier, the Treasury-Federal Reserve Accord of 1951, and, finally, the end of the deflationary policy to combat inflation and expectations of high inflation, and the beginning of the Great Moderation (Kim and Nelson, 1999; Perez-Quiros and McConnell, 2000; Gambetti and Galí, 2009; Canova, 2009).

The results are depicted in the scatter plots 2(b) to 2(e). The dotted line in each plot is our measure of the low-frequency relationship between inflation and deficits over debt. The slope of the scatter plots is equal to the ordinary least squares estimate of the following regression

$$\pi_t(\beta) = const + b_f d_t(\beta) + \text{error}, \tag{1}$$

where we assume orthogonality between $d_t(\beta)$ and the error term. The scatter plots point to the following characteristics of the low-frequency relationship. First, the co-movement changes over time, i.e. it is different in each sub-sample. Second, the slope of the dotted line is especially steep in the 1952–1983 sub-sample (Figure 2(d)). In the 1984–2009 sub-sample (Figure 2(e)) the dotted line is almost flat.

In order to accommodate the time-variation of the relationship, potential omitted variables, and in order to allow for lagged inflation in the estimation, we estimate the slope coefficient using a TVP-VAR model. Since the TVP-VAR model contains unfiltered data, we follow Sargent and Surico (2011) and make use of the finding by Whiteman (1984) and calculate the low-frequency relationship via the spectrum. We describe the setup of the TVP-VAR model in Section 3.2. In this section, we also provide more details on the estimation of the coefficient b_f, our measure of the low-frequency relationship.

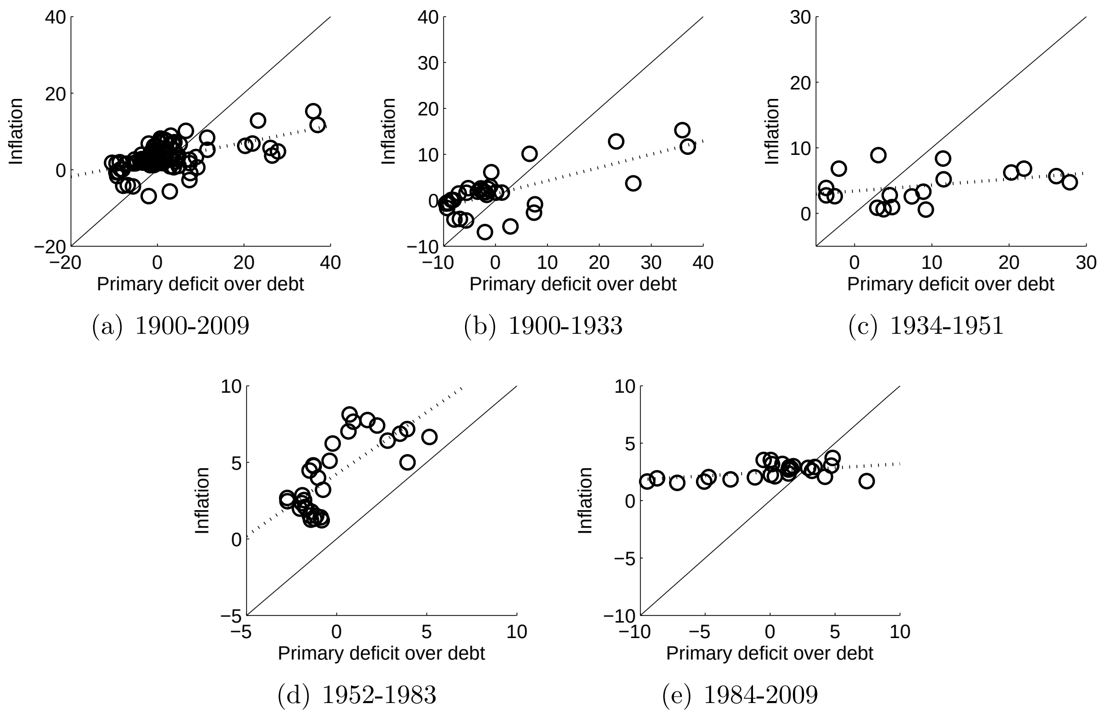


Figure 2: Scatter plots of filtered time series ($\beta = 0.95$) of inflation and deficits over debt. The dashed line indicates the slope of the scatter and the solid line is the 45° line.

3 The TVP-VAR model

In this section, we describe the data employed in the estimation, set up the TVP-VAR model, and specify the prior distribution for the coefficients.

3.1 Data

Next to inflation (π_t) and primary deficits over debt (d_t), we consider real output growth (Δx_t), money growth (Δm_t), and nominal interest rates (R_t) as additional observable variables. All variables are of quarterly frequency and range from 1875Q1 until 2011Q4. Real output growth is defined as year-to-year first differences of the logarithm of real GDP. From 1947Q1 onward, real GDP (in chained 2005 dollars) is taken from the FRED II database of the Federal Reserve Bank of St. Louis. For the period before 1947, we employ the growth rates of the real GNP series provided by Balke and Gordon (1986) to construct the time series. We apply the same procedure for money growth to the M2 stock series from the FRED II database starting in 1959Q1. For the period from 1875Q1 until 1981Q4, the nominal interest rate is defined as the six-month commercial paper rate given by Balke and Gordon (1986). From 1982Q1 onward, we use the quarterly averages of the weekly six-month Treasury Constant Maturity Rate available on the FRED II database. Finally, the sample spans the period 1876Q1 to 2011Q4.

3.2 Model setup

Given the vector of observable variables $\mathbf{y}_t = [d_t, \Delta x_t, \pi_t, R_t, \Delta M_t]$, the VAR model with time-varying coefficients and stochastic volatilities is defined as

$$\mathbf{y}_t = \mathbf{c}_t + \sum_{j=1}^{p} \mathbf{A}_{j,t}\mathbf{y}_{t-j} + \mathbf{u}_t = \mathbf{X}_t'\mathbf{A}_t + \mathbf{B}_t^{-1}\mathbf{H}_t^{\frac{1}{2}}\epsilon_t \ , \tag{2}$$

where $\mathbf{y}_t$ is a $n \times 1$ vector of macroeconomic time series, $\mathbf{c}_t$ is a time-varying $n \times 1$ vector of constants, $\mathbf{A}_{j,t}$ are p time-varying $n \times n$ coefficient matrices, and $\mathbf{u}_t$ is a $n \times 1$ vector of disturbances with time-varying variance-covariance matrix $\mathbf{\Omega}_t = \mathbf{B}_t^{-1}\mathbf{H}_t\left(\mathbf{B}_t^{-1}\right)'$. The time-varying matrices $\mathbf{H_t}$ and $\mathbf{B_t}$ are defined as

$$\mathbf{H_t} = \begin{bmatrix} h_{1,t} & 0 & \cdots & 0 \\ 0 & h_{2,t} & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \dots & 0 & h_{n,t} \end{bmatrix} \qquad \mathbf{B}_t = \begin{bmatrix} 1 & 0 & \cdots & 0 \\ b_{21,t} & 1 & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ b_{n1,t} & \dots & b_{n(n-1),t} & 1 \end{bmatrix}. \tag{3}$$

The time-varying coefficients are assumed to follow independent random walks with fixed variance-covariance matrices. In particular, laws of motions for the vector $\mathbf{a}_t = \text{vec}[\mathbf{c}_t\ \mathbf{A}_{1,t}\ ...\ \mathbf{A}_{p,t}]$, $\mathbf{h}_t = \text{diag}(\mathbf{H_t})$, and the vector $\mathbf{b}_t = [b_{21,t}, (b_{31,t}\ b_{32,t}), ..., (b_{n1,t}\ ...\ b_{n(n-1),t})]'$ containing the equation-wise stacked free parameters of $\mathbf{B}_t$ are given by

$$\mathbf{a}_t = \mathbf{a}_{t-1} + \nu_t, \tag{4}$$

$$\mathbf{b}_t = \mathbf{b}_{t-1} + \zeta_t, \tag{5}$$

$$\log \mathbf{h}_t = \log \mathbf{h}_{t-1} + \eta_t. \tag{6}$$

Finally, we assume that the variance-covariance matrix of the innovations is block diagonal:

$$\begin{bmatrix} \epsilon_t \\ \nu_t \\ \zeta_t \\ \eta_t \end{bmatrix} \sim N(0, V) \ , \text{ with } \mathbf{V} = \begin{bmatrix} \mathbf{I}_n & 0 & 0 & 0 \\ 0 & \mathbf{Q} & 0 & 0 \\ 0 & 0 & \mathbf{S} & 0 \\ 0 & 0 & 0 & \mathbf{W} \end{bmatrix} \text{ and } \mathbf{W} = \begin{bmatrix} \sigma_1^2 & 0 & \cdots & 0 \\ 0 & \sigma_2^2 & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \dots & 0 & \sigma_n^2 \end{bmatrix}, \tag{7}$$

where $\mathbf{I}_n$ is an n-dimensional identity matrix and $\mathbf{Q}$, $\mathbf{S}$, and $\mathbf{W}$ are positive definite matrices.

Moreover, it is assumed that matrix **S** is also block-diagonal with respect to the parameter blocks for each equation and **W** is diagonal.[2]

3.3 Prior specification

For the prior specification of the aforementioned model we follow the recent literature. In this regard, some of the prior parameters are based on a training sample from the beginning of the observation period (1876Q1–1899Q1). Therefore, we estimate a time-invariant VAR(2) model with ordinary least squares (OLS) and use the point estimates to calibrate some of the prior distributions (see, e.g., Cogley and Sargent, 2005; Primiceri, 2005). In particular, we use multivariate normal distributions for the priors on the initial conditions of the time-varying coefficients and the stochastic volatilities, which are parameterized with the corresponding OLS estimates. In addition, we use an inverse Wishart distribution for priors on the variance-covariance matrices of the error terms in the time-varying parameter equations, **Q** and **S**. For the prior on the diagonal elements of **W**, we use an inverse gamma distribution. The prior choices can be summarized as follows:

$$\begin{aligned}
\mathbf{a}_0 &\sim N(\hat{a}^{OLS}, 4\ Var(\hat{a}^{OLS})) \\
\mathbf{b}_0 &\sim N(\hat{b}^{OLS}, k_b^2\ I_4) \\
\log h_0 &\sim N(\log h^{OLS}, 10\ I_n) \\
\mathbf{Q} &\sim IW(k_Q^2\ 93\ Var(\hat{a}^{OLS}),\ 93) \\
\sigma_i^2 &\sim IG(\frac{k_w^2 \cdot 6}{2},\ \frac{6}{2}) \\
\mathbf{S}_1 &\sim IW(k_S^2\ 2\ I_1,\ 2) \\
\mathbf{S}_2 &\sim IW(k_S^2\ 3\ I_2,\ 3) \\
\mathbf{S}_3 &\sim IW(k_S^2\ 4\ I_3,\ 4) \\
\mathbf{S}_4 &\sim IW(k_S^2\ 5\ I_4,\ 5),
\end{aligned} \tag{8}$$

where the scaling factors k_Q and k_W are set to 0.01 and k_b and k_S to 0.1.

4 Estimation results

For the estimation of the model, we choose a lag length $p = 2$ and employ a Metropolis-within-Gibbs sampling algorithm. For a more detailed discussion of the estimation procedure, we refer the reader to Primiceri (2005) and Cogley and Sargent (2005). However, in

[2]See Primiceri (2005) for a discussion about relaxing these assumptions.

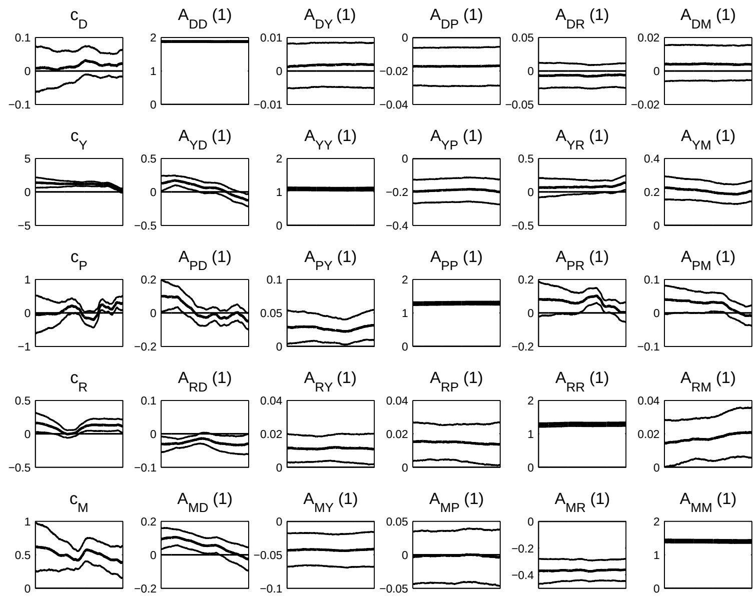
contrast to the former papers, we use a multi-move sampler to sample stochastic volatility as suggested by Shephard and Pitt (1997) and modified by Watanabe and Omori (2004).[3] During the simulation, we ensure stationarity of the VAR coefficients in the posterior distribution. Finally, we take 100,000 draws with a burn-in phase of 80,000 draws. We check for convergence by calculating various statistics and diagnostics which can be found in the corresponding appendix. After the burn-in phase, we keep only each 10th draw to reduce autocorrelation. This yields a sample of 2000 draws from the posterior density which is the basis for all presented results.

4.1 Parameter estimates

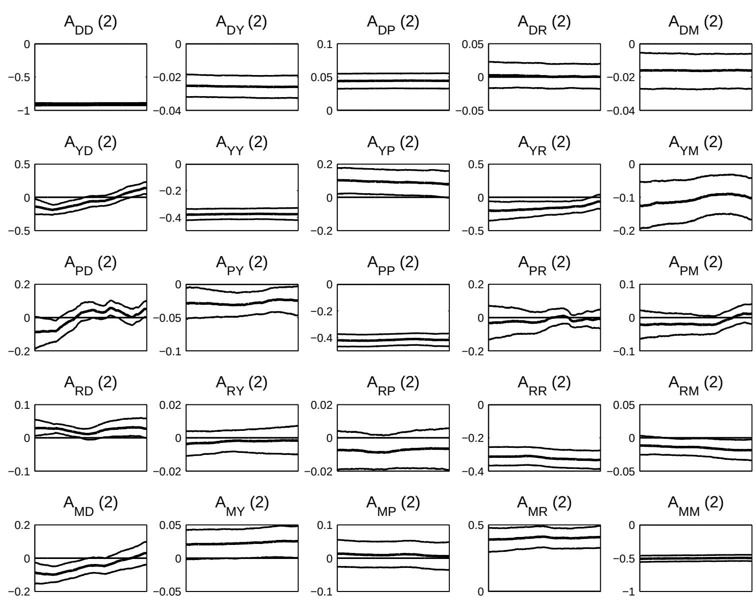
To shed light on the different sources of time variation in our TVP-VAR, we examine the time paths of the reduced-form coefficient matrix $\mathbf{A}_t$, the coefficients of the impact matrix $\mathbf{B}_t$, and the stochastic volatilities in $\mathbf{H}_t$. We begin our analysis with the change in the reduced form coefficients matrix $\mathbf{A}_t$. Figure 3 illustrates the time paths of the elements in $\mathbf{A}_t$, revealing that the degree of time variation varies considerably across coefficients. While, for instance, the VAR coefficients on own lags are almost constant, the coefficients that link the first lag of deficits over debt, interest rates, or money growth to inflation change substantially over the sample period. To scrutinize this heterogeneity in $\mathbf{A}_t$, we follow Cogley and Sargent (2005) and calculate the principal components of the variance-covariance matrix $\mathbf{Q}$ of the disturbances in the law of motion of the time-varying VAR coefficients. Similar to Cogley and Sargent (2005), we find that the change in $\mathbf{A}_t$ is driven by a few common components, indicating that most linear combinations of VAR coefficients are approximately constant. For the 55 free parameters in our VAR model, we find that only six principal components account for nearly all the variation. For the 10 free parameters of the time-varying contemporaneous matrix $\mathbf{B}_t$, we find a different picture. According to Figure 4, all coefficients vary considerably over time. The results from the principal component analysis confirm this observation. Although we have far fewer free parameters in $\mathbf{B}_t$ than in $\mathbf{A}_t$, we need also six principal components to account for more than 99 percent of the total variation.

Figure 5 shows the evolution of the square roots of the stochastic volatility of deficits over debt, real GDP growth, inflation, interest rates, and money growth. We find that our estimates of the stochastic volatilities coincide with those calculated in Sargent and Surico (2011). For real GDP growth, we observe that the Great Depression, followed by the two world wars, represents the most volatile period during the 20th century. After World War II and until the beginning of the Great Recession in 2007, we see an overall decrease in the

[3] See also Nakajima (2011) for a comparison and detailed description of this sampling algorithm.



(a) Constant and AR(1) parameter



(b) AR(2) parameter

Figure 3: Time-varying parameter estimates of matrix $\mathbf{A}$.

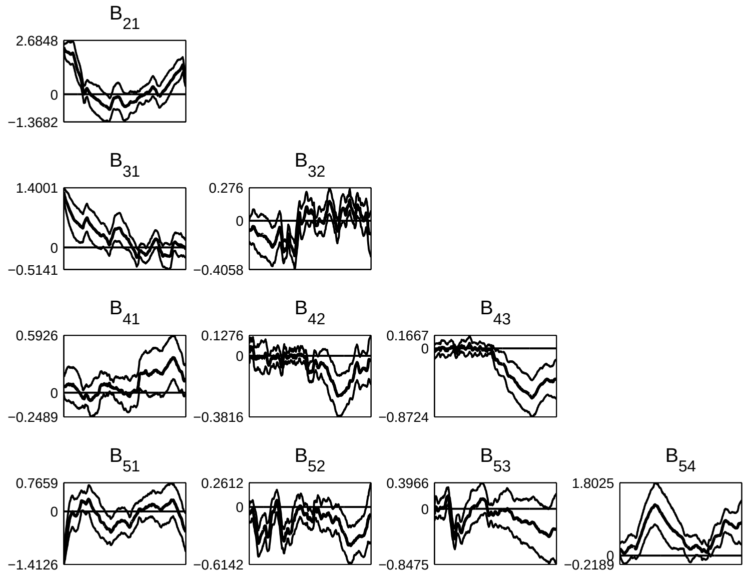
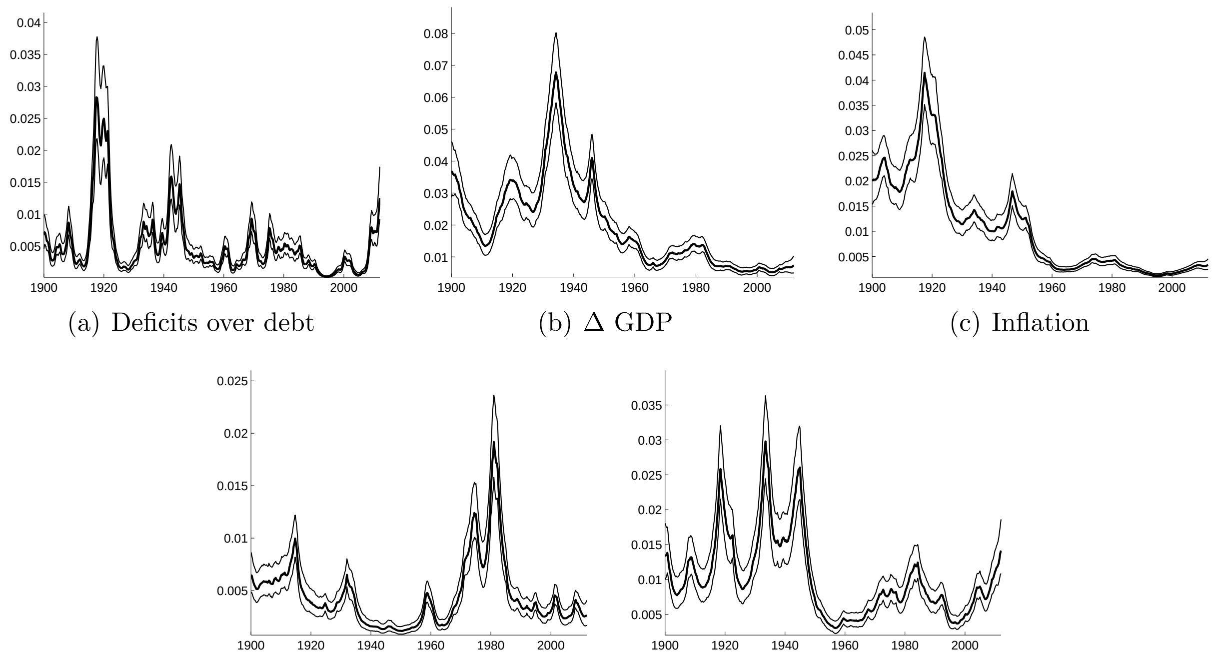


Figure 4: Time-varying parameter estimates of matrix **B**

magnitude of shocks to GDP growth, with the 1970s and the early 1980s representing the only exception. This overall decline in volatility during the post-World War II period is in line with the evidence in, e.g., DeLong and Summers (1986), where it is observed that postwar U.S. business cycles had become less volatile compared to the period before World War I, and certainly compared to the inter-war period. We also observe that shocks hitting the U.S. economy have been muted from the beginning of the 1980s up until the recent crises. This decline in the volatility of shocks is coined as the Great Moderation (see, e.g. Kim and Nelson (1999), Gambetti and Galí (2009), Canova (2009)). For the stochastic volatility component of inflation, the aftermath of World War I and the years of World War II seem to constitute the most volatile periods. Our results also indicate a high volatility under the Gold Standard and a relatively calm post-World War II period. Compared to the magnitude of the shocks during, e.g., World War I, even the apparently huge disturbances during the 1970s appear quite moderate. The estimates for the stochastic volatility of money growth reveal the usual suspects. The three most striking spikes occur after World War I, during the Great Depression, and around World War II. Moreover, it is remarkable that the size of shocks to the money growth equation in the VAR show similarly high values today to those last seen in the 1970s, indicating the consequences of the unconventional monetary policy actions since the Great Recession. The findings are different for shocks to the interest rate equation and the deficits over debt equation. For deficits over debt, we observe recurring episodes of shocks with high volatility. A possible explanation for this pattern lies in different episodes of fiscal intervention during and in the aftermath of various crises, such as the excessive war finance of World War I. Interestingly, the recent upsurge in U.S. deficit spending marks, following the two world wars, one of the three outstanding episodes in our sample. The results for the stochastic volatility of the interest rate equation pose another exception. In contrast to all other series, the biggest spikes in the volatility component are clustered between the end of the 1970s and the beginning of the 1980s, reflecting the anti-inflationary policy during this turbulent period.



(a) Deficits over debt (b) Δ GDP (c) Inflation

(d) 6m interest rate (e) Δ Money

Figure 5: Square roots of stochastic volatility.

4.2 The low-frequency relationship

Subsequently, we follow Sargent and Surico (2011) and make use of one result provided by Whiteman (1984). In particular, Whiteman (1984) shows that the low-frequency regression coefficient in equation (1) is an estimator of the sum of distributed lag coefficients, $\sum_{j=-\infty}^{\infty} \iota_j$, of the two-sided infinite least-squares projection of π_t on past, present, and future d's. Since this sum of lagged regression coefficients is equal to the cross spectrum of π and d, $S_{\pi d}$, divided by the spectrum of d, S_d, at frequency zero, the low-frequency relationship can be approximated via the spectral density by

$$b_f \approx \sum_{j=-\infty}^{\infty} \iota_j = \frac{S_{\pi d}(0)}{S_d(0)} \ . \tag{9}$$

To approximate the evolution of this parameter over time, we calculate the temporary spectral density from the estimates of the TVP-VAR model. To this end, we write the TVP-VAR model at time t conditional on T in state-space form

$$\begin{aligned} \mathbf{X}_t &= \hat{\mathbf{A}}_{t|T}\mathbf{X}_{t-1} + \hat{\mathbf{B}}_{t|T}\mathbf{w}_t \\ \mathbf{y}_t &= \hat{\mathbf{C}}_{t|T}\mathbf{X}_t \ , \end{aligned} \tag{10}$$

where $\mathbf{X}_t$ is the $n_x \times 1$ state vector, $\mathbf{y}_t$ is an $n_y \times 1$ vector of observables, $\mathbf{w}_t$ is an $n_w \times 1$ Gaussian random vector with mean zero and unit covariance matrix that is distributed identically and independently across time. The matrices $\hat{\mathbf{A}}$, $\hat{\mathbf{B}}$, and $\hat{\mathbf{C}}$ are functions of a vector of the time-varying structural model parameters. Given this representation, the corresponding temporary spectral density at time $\mathbf{t}$ of matrix Y is

$$S_{Y,t|T}(\omega) = \hat{\mathbf{C}}_{t|T}\left(I - \hat{\mathbf{A}}_{t|T}e^{-i\omega}\right)^{-1}\hat{\mathbf{B}}_{t|T}\hat{\mathbf{B}}'_{t|T}\left(I - \hat{\mathbf{A}}'_{t|T}e^{i\omega}\right)^{-1}\hat{\mathbf{C}}'_{t|T} \tag{11}$$

and the temporary low-frequency relationship between deficits over debt and inflation at time t is computed as

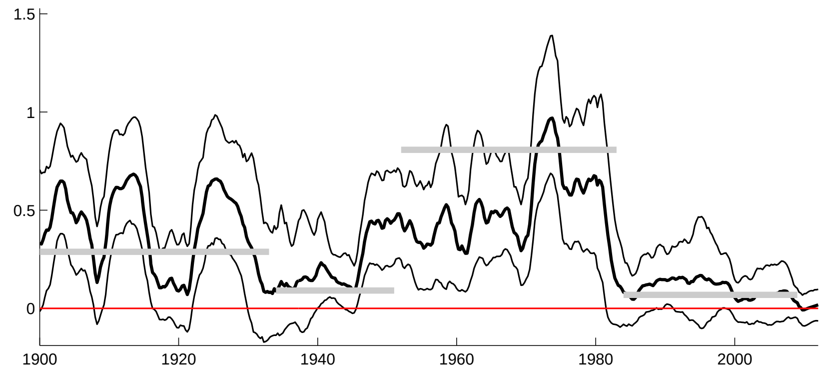
$$\hat{b}_{f,t|T} = \frac{S_{\pi,d,t|T}(0)}{S_{d,t|T}(0)} \tag{12}$$

Moreover, the corresponding measure of fit which can be interpreted as R^2 is defined as coherency squared:

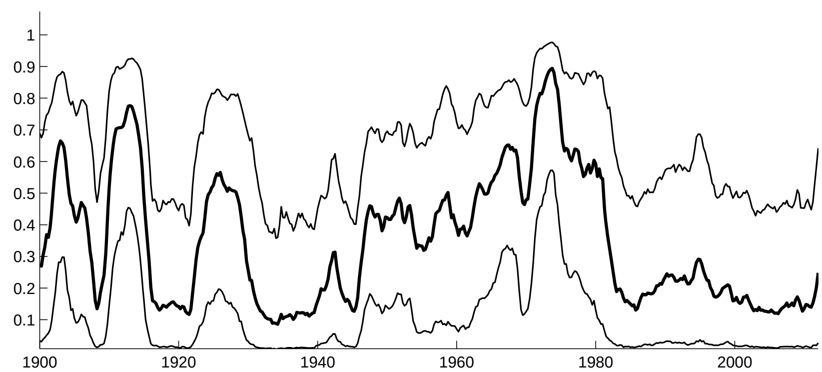
$$\Upsilon_{\pi d,t|T}(0) = \frac{|S_{\pi d,t|T}(0)|^2}{S_{\pi,t|T}(0)S_{d,t|T}(0)} \tag{13}$$

When we estimate the temporary object in (12), we do not account for the fact that the parameter drift is going forward beyond time t. This simplification is necessary because integrating out the high dimensional predictive density for all possible paths is computationally not feasible. Following Sargent and Surico (2011), we relate the assumption to the "anticipated utility" approach in the learning literature. In this circumstance, we assume that the agents do not know the true model and treat the actual values in period t such as they remain forever. However, this assumption could be problematic in the case of frequent non-persistent changes to the parameters. Nevertheless, as pointed out by Kreps (1998), such behavior by the agents constitutes a rational action, in which the agents have difficulty identifying the true model.

The results of the TVP-VAR model are illustrated in Figure 6, where sub-plots show the evolution of the low-frequency relationship between inflation and the primary deficit-to-debt ratio (Figure 6(a)) and coherency squared as the corresponding measure of fit (Figure 6(b)). Figure 6(a) illustrates that the time-variation of the low-frequency relationship is in line with the different slopes of the scatter plots which are based on the OLS regression coefficient of the filtered data in Section 2. Hence, we conclude that the spectrum is well approximated by the TVP-VAR model.



(a) $\hat{b}_f$: Median and 68% central posterior bands for the time-varying regression coefficient of inflation on deficits over debt. Grey lines depict slopes of the scatter plots which are based on the OLS regression coefficient of the filtered data in Section 2.



(b) R^2: Inflation on deficits over debt

Figure 6: Median and 68% central posterior bands for $\hat{b}_f$ and corresponding R^2

Our first main finding is that the low-frequency relationship between the variables of interest is time-varying. Except for a short period after World War I and the times around the Great Depression, this relationship is significantly positive until 1980. While the relationship in the first part of the 20th century is rather volatile, it is stable in the period from 1945 to the end of the Bretton Woods system. In the year the Bretton Woods system ends, the low-frequency relationship increases to its highest value, around one. It decreases sharply in 1979 when Paul Volcker became Chairman of the Federal Reserve System. After 1980, zero is included in the probability band most of the time. This period from 1960 to 1980 is also the time of the highest coherence.

To investigate the sensitivity of our main finding, we conduct various robustness checks. First, while interpolating the annual time series of government debt held by the public allows us to include more observations and to obtain more precise estimates, it introduces the possibility of an approximation error. Hence, we check the robustness of our estimation by considering two different interpolation techniques. Second, the long time series for government debt is only available in par values and not in market values. The construction of a market-value government debt series, which obeys the flow budget constraints, requirers various adjustments. This comes at the cost of potential measurement error. For example, Chung and Leeper (2007) point out that such a constructed debt series is not generally consistent with other time series, such as, Cox and Hirschhorn (1983). Nevertheless, we analyze the robustness of our results with respect to interpolation and market value of debt by constructing a quarterly proxy, with quarterly primary deficits calculated from NIPA data and market value of privately held gross federal debt calculated by the Federal Reserve Bank of Dallas. The final quarterly time series covers the time from 1947Q1 until 2010Q1. Third, we also perform a robustness exercise regarding the selected variables and their ordering in the VAR model. To this end, we substitute the measure of fiscal stance by considering debt growth instead of deficits over debt. Moreover, we run robustness checks by using real as well as nominal 3-month interest rates instead of the 6-month commercial paper rate and by substituting the GDP deflator with the CPI deflator, which is the time series "Consumer Price Index for All Urban Consumers (CPIAUCSL)" taken from the FRED II database of the Federal Reserve Bank of St. Louis. Since, the latter series is available only from 1947Q1 onward, we run the estimation with a shorter sample. Finally, while this robustness check involves a different training sample for the prior calibration, we build up confidence that our results are also robust in this respect. For completeness, we also analyze an alternative approach to estimate the low-frequency relationship directly from unfiltered data, namely the dynamic OLS estimator as proposed by Stock and Watson (1993). Moreover, we investigate miscellaneous time-invariant BVAR models with different lag specifications. To conclude, all aforementioned robustness checks satisfactorily confirm our finding.[4]

4.3 Structural decomposition of the low-frequency relationship

In our structural decomposition of the low-frequency relationship, we identify a monetary policy shock and investigate how much of the variation in the low-frequency relationship is due to the long-run effects of the monetary policy shock. The monetary instrument we consider is interest rates, because we are especially interested in the period from 1970 onwards at this stage. The monetary policy shock is identified by a recursiveness assumption. In particular, we assume that shocks to the monetary instrument do not affect variables

[4] All robustness checks and other additional material can be found in the corresponding online appendix, e.g., on one of the authors' websites: `http://www.mwpweb.eu/AlexanderKriwoluzky/`

ordered above the monetary instrument in the first period. The variables are ordered as in the model setup described in Section 3.2, i.e. deficits over debt, output growth, inflation, interest rates, and money growth. By ordering the fiscal variable first, we assume its unresponsiveness to other shocks in the first quarter, which is in line with the assumption made by Blanchard and Perotti (2002). Moreover, the ordering of the remaining variables is taken from the related literature (see, e.g, Sims, 1992; Christiano et al., 1996), where it is typically assumed that output growth and inflation respond with a one-period lag to a monetary policy shock, while the monetary aggregate responds immediately. By applying the aforementioned identification scheme, we use the Cholesky decomposition of the reduced form variance-covariance matrix, $\mathbf{\Omega}$, and denote the corresponding expanded matrix in the state-space system as $\tilde{\mathbf{B}}$. Accordingly, the re-written state-space system is given by

$$\begin{aligned} \mathbf{X}_t &= \hat{\mathbf{A}}_{t|T}\mathbf{X}_{t-1} + \tilde{\mathbf{B}}_{t|T}\epsilon_t \\ \mathbf{y}_t &= \hat{\mathbf{C}}_{t|T}\mathbf{X}_t \ . \end{aligned} \tag{14}$$

Furthermore, denote the i-th column in the matrix $\tilde{\mathbf{B}}$ by $\tilde{\mathbf{B}}^{\mathbf{i}}$. The temporary spectrum conditional on the i-th structural shock at time t is given by:

$$S^i_{Y,t|T}(\omega) = \hat{\mathbf{C}}_{t|T}\left(I - \hat{\mathbf{A}}_{t|T}e^{-i\omega}\right)^{-1}\tilde{\mathbf{B}}^i_{t|T}(\tilde{\mathbf{B}}^i)'_{t|T}\left(I - \hat{\mathbf{A}}'_{t|T}e^{i\omega}\right)^{-1}\hat{\mathbf{C}}'_{t|T} \tag{15}$$

Since the structural shocks are independent, we know that the unconditional spectrum in equation (15) is the sum of spectra conditional on each structural shock.[5] Consequently, the unconditional low-frequency measure can be written as the sum of weighted conditional low-frequency measures

$$\hat{b}_{f,t|T} = \frac{S^m_{d,t|T}(0)}{S_{d,t|T}(0)}\hat{b}^m_{f,t|T} + \sum_{i=1}^{4}\frac{S^i_{d,t|T}(0)}{S_{d,t|T}(0)}\hat{b}^i_{f,t|T} \ , \tag{16}$$

where the low-frequency relationship conditional on the monetary policy shock is given by $\hat{b}^m_{f,t|T}$. The weight in front of this conditional low-frequency measure, $\frac{S^m_{d,t|T}(0)}{S_{d,t|T}(0)}$, is the fraction of the unconditional spectrum at frequency zero explained by the corresponding spectrum conditional on the monetary policy shock. Since we are mainly interested on the long-run effects of monetary policy, we simply add up the remaining weighted low-frequency measures conditional on non-monetary policy shocks. Subsequently, we present the median of the time-varying unconditional objective decomposed into the weighted measures conditional on

[5]For similar approaches see, e.g., Gambetti and Galí (2009), Gambetti, Pappa, and Canova (2008), and Mertens (2010).

monetary policy and conditional on non-monetary policy shocks. Figure 7 shows the main result: while the monetary policy shock contributes only to a small extend to the movements of the low-frequency relationship before 1960, most parts of the increase in the low-frequency relationship during the 1960s and 1970s can be attributed to the effects of a monetary policy shock.

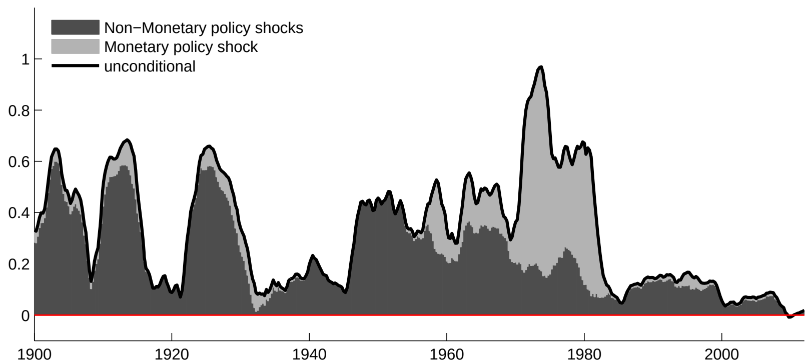


Figure 7: Historical decomposition of the unconditional low-frequency relationship between deficits over debt and inflation.

In order to gain further insights into the importance of the transmission mechanism of monetary policy shocks, we perform a counterfactual analysis. In this regard, we investigate whether the movements of the low-frequency relationship are due to changes in the volatilities of the shocks or whether the changes in the low-frequency relationship can be attributed to changes in the systematic behavior of the economy. To that end, we fix the systematic behavior of the economy (the matrices $\mathbf{A}_t$ and $\mathbf{B}_t$ of the VAR model in eq. (2)) in a first counterfactual experiment to the first quarter of 1995 and in a second experiment to the first quarter of 1976.[6] More precisely, for the first experiment, we fix the systematic behavior of the economy to be $\mathbf{A}_{1995.1}$ and $\mathbf{B}_{1995.1}$ at each point in time, i.e. we draw realizations for $\mathbf{A}_{1995.1}$ and $\mathbf{B}_{1995.1}$ out of their posterior distributions. For every draw, the matrix ($\mathbf{H}_t$) is drawn from its posterior distribution at each point in time and we calculate the unconditional low-frequency relationship using equation (12). The second experiment repeats the exercise, but with fixed matrices $\mathbf{A}_{1976.1}$ and $\mathbf{B}_{1976.1}$. Figures 8 and 9 display the unconditional low-frequency relationship and its decomposition into monetary and non-monetary policy shocks of the counterfactual experiments.

Comparing Figures 8 and 9 with Figure 7, we first deduce that the movements of the low-frequency relationship in the first part of the 20th century are mostly attributable to the changes in the volatilities of the shocks in the economy. Furthermore, the figures indicate that

[6]In the corresponding online appendix, we also conduct robustness exercises with respect to different dates and averages over periods. Moreover, we present a counterfactual with respect to the size of the shocks by fixing $\mathbf{H}_t$ over time. All these exercises confirm the findings in this section.

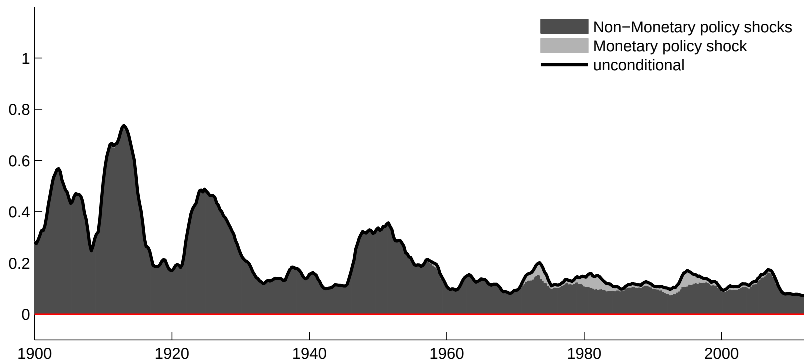


Figure 8: Counterfactual experiment: Median of $\hat{b}_f$ for fixed VAR model coefficients ($\mathbf{A}_{1995.1}$ and $\mathbf{B}_{1995.1}$).

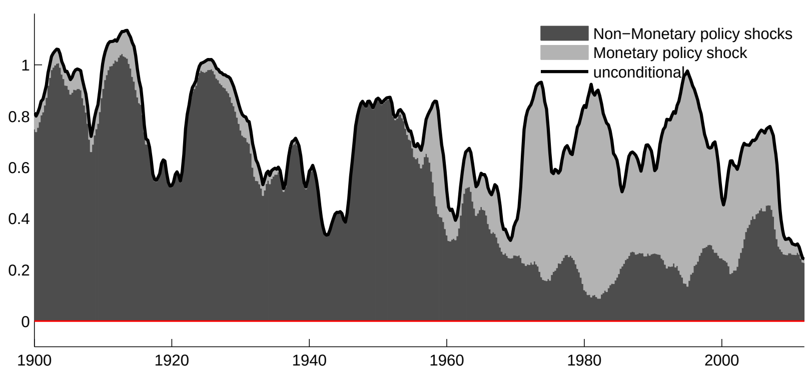


Figure 9: Counterfactual experiment: Median of $\hat{b}_f$ for fixed VAR model coefficients ($\mathbf{A}_{1976.1}$ and $\mathbf{B}_{1976.1}$).

the decline in the low-frequency relationship in 1980 is due to a change in the transmission mechanism of monetary policy shocks. Had the economy been in the same state it was in 1976Q1 all the time, the low-frequency relationship after 1980 would not have been around zero – and the long-run impact of monetary policy shocks would have been the main driving force. On the contrary, had the economy been in its 1995Q1 state the whole time, the high low-frequency relationship in the 1960s and 1970s would not have been estimated. This result is closely related to the findings of Bianchi and Ilut (2012). Therefore, we interpret our results along their finding that a change in the interaction of the monetary and fiscal policy was responsible for the change in the transmission mechanism of a monetary policy shock. In the following, we provide additional anecdotal evidence and relate our result to economic theory.

Our results are related to broad anecdotal evidence. For example, Meltzer (2010) characterizes the period of the 1960s and 1970s as one of the Fed accepting "its role as a junior partner by agreeing to coordinate actions with the administration's fiscal policy." Similarly, Greider (1987) argues that Arthur Burns ran an unusually expansionary policy because he believed it would increase his chances of being nominated for another term. However, the strong low-frequency relationship declines sharply after 1980, i.e. after Paul Volcker became Fed Chairman. As Meltzer (2010) points out, Volcker rebuilt much of the independence and credibility the Federal Reserve had lost during the two previous decades. In this regard, Martin (2013) presents the number of meetings at the White House between the U.S. President and the Fed Chairman. He shows that the number of meetings with Presidents Nixon and Ford (1969-1977) were quite frequent and took place four times more often than the next four presidents put together. Additionally, Martin (2013) shows that President Johnson (1963-1969) met with the Fed Chairman 300 times during his five years in office.

While the monetary authority started to become independent and started to fight inflation under the chairmanship of Paul Volcker in the beginning of the 1980s, fiscal policy conducted by the Reagan Administration cannot be necessarily seen as backing the outstanding debt by future primary surpluses and thus accommodating the actions of the monetary authority. In the following, we employ narrative evidence obtained from Romer and Romer (2010) and argue that what matters are the expectations of the public. More precisely, it matters whether the public expected the fiscal authority to back the government debt eventually by raising taxes or decreasing government expenditures. While the Reagan Administration enacted the Economic Recovery Tax Act of 1981, the following tax changes were all aimed at reducing the deficits: the Tax Equity and Fiscal Responsibility Act of 1982, the Social Security Amendments of 1983, and finally the Deficit Reduction Act of 1984. The present value of the latter tax changes (in total \$70.15 billion) has been smaller than the present value of the tax reduction act ($125.90 billion), but we interpret these tax changes as a signal to the public that the fiscal authority is concerned about the state of fiscal finances. Moreover, the remaining years of the 1980s are characterized by the Omnibus Budget Reconciliation Act of 1987 and of 1990, again meant to reduce the fiscal deficit. In terms of their present value, these Budget Reconciliation Acts are larger than the present value of the tax reduction by the Tax Reform Act of 1986 ($45.14 billion vs. $10.12 billion). Thus, we interpret this narrative evidence as suggesting that the public expected the fiscal authority to accommodate the actions by the central bank.

Regarding economic theory, our results are also in line with the theory about monetary and fiscal interaction. The FTPL suggests that if the fiscal authority does not back current outstanding debt by future primary surpluses and the monetary authority is not applying the Taylor principle, a contractionary monetary policy shock actually leads to an increase in inflation and has prolonged and long-lasting effects on inflation and public debt. These effects are stressed by Sims (2011) who analyzes the situation of the Federal Reserve Board in the 1970s and concludes that even, if the central bank aims to fight inflation by increasing the nominal interest rate, it actually increases inflation in the long run. Sims (2011) labels his result as the central bank "stepping on a rake". Contrarily, in a situation of monetary dominance, where fiscal authority raises sufficiently high primary surpluses to serve outstanding debt and the monetary authority fulfills the Taylor principle, monetary policy shocks have no prolonged effects. Consequently, the breakdown of the low-frequency relationship between deficits over debt and inflation after 1980 can be interpreted as a change in the interaction of monetary and fiscal policy.

Similarly, the unpleasant monetarist arithmetic suggests, for times where fiscal policy is unable or unwilling to back current outstanding debt by future primary surpluses, that the central bank can step in by providing seignorage, which eventually leads to inflation. Thus, inflation and deficits over debt are highly correlated at low frequencies. On the other hand, if monetary policy dominates and the central bank is independent, there is no relationship between our variables of interest. Therefore, we interpret our findings as confirmation that the low-frequency relationship between deficits over debt and inflation is strongly related to the independence of monetary policy and its interaction with fiscal policy.

5 Conclusion

The low-frequency relationship between inflation and primary deficits over lagged debt for the U.S. is time-varying and mostly positive between 1900 and 1980. We find the strongest relationship neither in times of crisis nor in times of high public deficits, but from the mid-

1960s up to 1980. After 1980 it becomes insignificantly different from zero. Our result shows that the low-frequency relationship between fiscal deficits and inflation is strongly related to the conduct of monetary policy and its interaction with fiscal policy after World War II. In particular, we find that most of the relationship in the 1960s and 1970s, as well as the sharp drop in 1980, are related to a change in the transmission mechanism of monetary policy shocks. This finding confirms the literature that the relationship between fiscal deficits and inflation depends on the monetary/fiscal policy mix in place.

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