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**Working Paper**

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# Monetary-Fiscal Policy Interaction and Fiscal Inflation: A Tale of Three Countries*

## Abstract

We study the impact of the interaction between fiscal and monetary policy on the low-frequency relationship between the fiscal stance and inflation using cross-country data from 1965 to 1999. In a first step, we contrast the monetary-fiscal narrative for Germany, the U.S. and Italy with evidence obtained from simple regression models and a time-varying VAR. We find that the low-frequency relationship between the fiscal stance and inflation is low during periods of an independent central bank and responsible fiscal policy and more pronounced in times of high fiscal budget deficits and accommodative monetary authorities. In a second step, we use an estimated DSGE model to interpret the low-frequency measure structurally and to illustrate the mechanisms through which fiscal actions affect inflation in the long run. The findings from the DSGE model suggest that switches in the monetary-fiscal policy interaction and accompanying variations in the propagation of structural shocks can well account for changes in the low-frequency relationship between the fiscal stance and inflation.

Keywords: time-varying VAR, inflation, public deficits

JEL Classification: E42, E58, E61

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# 1 Introduction

In a recent article, Summers (2014) paints a dire picture of future macroeconomic developments, pointing to the risk of a secular stagnation with a long period of poor economic growth and permanent negative natural rates of interest. He highlights policy interventions that could further reduce the real interest rate as a possible way out. With nominal interest rates close to zero, monetary policy alone is not up to the task, calling for fiscal policy to step in to reduce real interest rates through subdued fiscal inflation. Economic models that allow fiscal policy to play an important role in determining the price level - such as the fiscal theory of the price level (FTPL) - are thus brought to the center of attention.[1] However, despite their increasing importance in the policy debate the evidence in favor of the FTPL still remains scarce and is mostly limited to the experience of the U.S.[2]

In this paper, we provide further cross-country evidence for the FTPL by contrasting the experience of the U.S. between 1965 and 1999 with Italy and Germany, which are well known to have had different monetary and fiscal policy interactions in place.[3] We focus on the relationship between the fiscal stance and inflation and proceed in two complementary ways. We first estimate the time-varying low-frequency relationship between these two variables using a medium-sized time-varying VAR model (TVP-VAR). The focus on the low frequency is in the spirit of Lucas (1980) who suggests that change in the systematic relationship between two variables is best recovered beyond business cycle frequencies. In a second step, we employ a DSGE model to interpret the low-frequency measure structurally and to illustrate the mechanisms through which fiscal actions affect inflation in the long run. In this exercise, we allow specifically for a switch in the policy regime, i.e. the interaction between monetary and fiscal policies.

Our findings from the time-varying VAR suggest that the low-frequency relationship between the fiscal stance and inflation varies across time and country. More importantly, the evolution of the low-frequency relationship is strikingly consistent with the narrative evidence on the interaction of the monetary and fiscal authority for all three countries. For Italy, we find a high low-frequency relationship up to the late 1980s and a pronounced drop in the relationship in the early 1990s. This empirical result corresponds to the fact that Banca d'Italia was required by law to buy government securities at a fixed interest rate during the 1970s, moved gradually towards independence in the early 1980s, and became independent in

[1] The first studies to develop a theory for the interaction between monetary and fiscal policies are Sargent and Wallace (1981) and Leeper (1991).

[2] The notable exception is the study by Loyo (2000) who considers Brazil in the 1980s.

[3] Our choice of the set of countries is in line with Brunner, Fratianni, Jordan, Meltzer, and Neumann (1973), who studied the influence of monetary and fiscal policies on inflation between 1948 and 1971.

the run-up to the Maastricht Treaty which Italy signed up to during the 1990s. For Germany, the low-frequency relationship fluctuates around zero throughout the sample. This virtually non-existent relationship corresponds to the well-established fact that, as early as the 1970s, Germany had an independent central bank focusing on price stability and a fiscal policy which backed outstanding government debt. For the U.S., the inauguration of Paul Volcker as Fed chair coincides with the biggest drop in the estimated low-frequency relationship between the fiscal stance and inflation.[4] This empirical finding indicates that U.S. monetary policy possibly accommodated fiscal policy during the pre-Volcker era and determined the inflation rate in combination with a fiscal authority that backed outstanding government debt after Paul Volcker became chairman of the Federal Reserve in 1979.[5]

Given the findings of the TVP-VAR model we investigate further whether the change in the low-frequency relationship is indeed due to a change in the interaction between monetary and fiscal policies. Using a counterfactual experiment of the TVP-VAR model, we first demonstrate that the change in the low-frequency relationship in Italy as well as in the U.S. cannot be attributed to a change in the volatilities of the structural shocks. Second, we estimate a standard DSGE model on U.S. data from 1984 to 2009. We fix the volatilities of the structural shocks to their estimated values and perform a prior predictive analysis for the remaining parameters. The prior predictive analysis reports the probability distribution of the low-frequency relationship that a particular policy regime can produce before taking the model to the data. The results of the prior predictive analysis pinpoints the fact that the policy regime is the crucial element of the DSGE model determining the low-frequency relationship between fiscal stance and inflation. In particular, a high low-frequency relationship between the fiscal stance and inflation - as observed in our empirical analysis - is much more likely within an FTPL model setup. Therefore, our findings confirm the aforementioned narrative evidence that a change in the fiscal and monetary policy interaction can explain the change in the low-frequency relationship of interest.

The paper is structured as follows: Section 2 describes the dataset, sets up the time series model, presents the evolution of the low-frequency relationship, and relates its evolution to narrative evidence of each country; Section 3 sets up the medium-scale DSGE model to show how the changes in the low-frequency relationship are related to changes in the interaction between monetary and fiscal policies; Section 4 concludes.

[4]See Kliem, Kriwoluzky, and Sarferaz (2015) for a more detailed discussion.

[5]See, for example, Bianchi (2012), Bianchi and Ilut (2014), and Chen, Leeper, and Leith (2015).

# 2 Three countries, inflation, and the fiscal stance

In this section, we describe the dataset, set up the time series model, estimate the low-frequency relationship between inflation and the fiscal stance, and relate the estimation results to the narrative accounts for Italy, Germany, and the U.S.

## 2.1 Data

In this subsection, we describe the data sources and the transformation of the data. Following Sims (2011) and our previous work (Kliem et al., 2015), we employ primary deficits over one-period lagged debt as a measure of fiscal stance. 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 – government's primary surplus over one-period lagged debt. This summarizes net payments to bondholders either through interest rates or through the retirement of bonds. In other words, it represents the decrease in the fiscal authority's future liabilities. Conversely, a change in the deficits over debt measures the change in the fiscal authority's future liabilities. 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.

For each country, we establish a data set which contains primary deficits over debt, inflation, real GDP growth, nominal interest rates, and money growth. To estimate the time variation of the low-frequency relationship between the variables as precisely as possible, we choose for each country the longest coherent time period available. For some variables, this decision has inherent data limitations, which we will describe in more detail. The finally employed time series range from 1876Q1 until 2011Q4 for the U.S. and range from 1961Q1 until 1998Q4 for Italy and Germany, respectively. The available time span for Germany and Italy is limited for the following reasons. First, there are no coherent time series available for the time before and during World War II. Second, the introduction of the euro in 1999 marks a natural end to the countries' individual monetary-fiscal policy mix.

The fiscal time series for primary deficits over debt ($d_t$) for each country is constructed as follows. For the U.S., we use the time series for primary deficit and government debt held by the public from Bohn (2008).[6] For Italy and Germany, we make use of the fiscal database provided by Mauro, Romeu, Binder, and Zaman (2013). Because, the fiscal

[6]See `http://www.econ.ucsb.edu/~bohn/morepapers.html` for more details and recent updates of these time series. A detailed description of the fiscal data set and its construction is given by Bohn (1991).

database contains only ratios relative to GDP, we use annual GDP data from the IMF IFS database to construct the deficit over debt variable. Moreover, all fiscal time series are of annual frequency only. Therefore, we decide to interpolate the annual data using the cubic-spline approach. Additionally, time series for government debt for Italy and Germany are available only in par values and not in market values, while market values for the U.S. are available only from 1942 onward. Since we are interested in the low-frequency relationship of the variables, temporary differences between market and par values are not critical (see also Bohn, 1991).[7]

The remaining variables for the U.S. are calculated as follows. Inflation ($\pi_t$) is measured as year-to-year first differences of the GDP deflator. Following Sargent and Surico (2011), we use the data taken from the FRED II database starting in 1947Q1 and from Balke and Gordon (1986) previously. Similarly, real output growth ($\Delta x_t$) is defined as year-to-year first differences of the logarithm of real GDP. From 1947Q1 onward, real GDP (in chained 2010 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 ($\Delta M_t$) to the M2 stock series from the FRED II database starting in 1959Q1. For the nominal interest rate ($R_t$), we use the quarterly average of the effective Federal Fed Funds rate from 1954Q3 onward extended by the short-term interest rate from Balke and Gordon (1986) for the time previously.

For Italy, inflation is measured as year-to-year first differences of the CPI deflator from 1960Q1 onward taken from the IMF IFS database. Real output growth is calculated as year-to-year first differences of the logarithm of real GDP available from the OECD Quarterly National Accounts. Money growth is calculated as year-to-year first differences of the logarithm of M2 stock available from the Banca d'Italia, which is seasonal adjusted using Census x13. For the nominal interest rate, we employ the IMF IFS database again. In particular, from 1977Q3 until 1998Q4 we use the provided Treasury Bill rate and extend the series with the interest rate on government securities for the preceding period.

Because of the reunification of Germany, the construction of a coherent data set needs some more adjustments to avoid jumps. In particular, we use nominal GDP data and the corresponding GDP deflator from 1991Q1 until 1998Q4, which are extended using corresponding growth rates for West Germany for the period from 1970Q1 until 1989Q4. All data are taken from the Bundesbank. From these series, we construct real GDP for 1970Q1 onward which is finally extended until 1960Q1 by using growth rates of real GDP provided

[7]For a detailed discussion and extensive robustness checks regarding interpolation and market value of debt in addition to other changes of specifications, see Kliem et al. (2015).

by the IMF IFS database. Similarly, we combine the CPI deflator for unified Germany from 1991Q1 onward with the CPI deflator for West Germany from 1960Q1 until 1990Q4, where both series are taken from IMF IFS database. Finally, we calculate our inflation measure as year-to-year first differences of the logarithm of this constructed series. As the time series for the nominal interest rate, we use the T-Bill rate series from the IMF IFS database from 1977Q3 until 1998Q4 and the Money market rate for the preceding time.

## 2.2 The time-varying parameter VAR model

For each country, we estimate a single TVP-VAR model with the vector of observable variables $\mathbf{y}_t = [d_t, \Delta x_t, \pi_t, R_t, \Delta M_t]$. Each VAR model with time-varying coefficients and stochastic volatilities is defined as

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

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

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

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

$$\mathbf{a}_t = \mathbf{a}_{t-1} + \nu_t, \tag{3}$$

$$\mathbf{b}_t = \mathbf{b}_{t-1} + \zeta_t, \tag{4}$$

$$\log \mathbf{h}_t = \log \mathbf{h}_{t-1} + \eta_t. \tag{5}$$

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

$$\begin{bmatrix} \epsilon_t \\ \nu_t \\ \zeta_t \\ \eta_t \end{bmatrix} \sim N(0, V) \text{ , with } \mathbf{V} = \begin{bmatrix} \mathbf{I}_n & 0 & 0 & 0 \\ 0 & \mathbf{Q} & 0 & 0 \\ 0 & 0 & \mathbf{S} & 0 \\ 0 & 0 & 0 & \mathbf{W} \end{bmatrix} \text{ and } \mathbf{W} = \begin{bmatrix} \sigma_1^2 & 0 & \cdots & 0 \\ 0 & \sigma_2^2 & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \dots & 0 & \sigma_n^2 \end{bmatrix} , \quad (6)$$

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 $\mathbf{S}$ is also block-diagonal with respect to the parameter blocks for each equation and $\mathbf{W}$ is diagonal.

For the prior specifications of the aforementioned models we follow the recent literature. In this regard, some of the prior parameters are based on a training sample with a length of 40 quarters from the beginning of each observation period. More precisely, 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 (e.g., Cogley and Sargent, 2005; Primiceri, 2005). Like Bianchi and Civelli (2015), we choose the same hyperparameters across the TVP-VAR models, however, given the training sample approach the final priors for each model are different. See Appendix B.1 for a more detailed description of our choice of priors.

For the estimation of each TVP-VAR model, we choose a lag length $p = 2$ and employ a Metropolis-within-Gibbs sampling algorithm as described in Kliem et al. (2015). During the simulation, we ensure stationarity of the VAR model coefficients in the posterior distribution. We take 250,000 draws with a burn-in phase of 230,000 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 2,000 draws from the posterior density, which is the basis for all results presented throughout the paper.

## 2.3 The low-frequency relationship

As our measure for the low-frequency relationship between two variables, we follow the suggestion by Lucas (1980). After filtering the data to extract the low-frequency components of each time series, Lucas (1980) computed the regression coefficient in an ordinary-least-square regression. In our case, the variables of interest are deficits over debt and inflation. We denote the regression coefficient of a regression of deficits on inflation by $b_f$. Whiteman

(1984) shows that the regression coefficient can be approximated in the following way:

$$b_f \approx \frac{S_{\pi d}(0)}{S_d(0)} \ , \tag{7}$$

where $S_d$ is spectrum of $d$ and $S_{\pi d}$ is the cross spectrum of $\pi$ and $d$ at frequency zero. In order to estimate the relationship using unfiltered data in the TVP-VAR model, we follow the procedure described in Sargent and Surico (2011). We use the state-space representation of the TVP-VAR model:

$$\begin{aligned} \mathbf{X}_t &= \hat{\mathbf{A}}_{t|T}\mathbf{X}_{t-1} + \hat{\mathbf{B}}_{t|T}\mathbf{w}_t \\ \mathbf{y}_t &= \hat{\mathbf{C}}_{t|T}\mathbf{X}_t \ , \end{aligned} \tag{8}$$

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. The corresponding spectral density at time $\mathbf{t}$ of matrix $Y$ hence is

$$S_{Y,t|T}(\omega) = \hat{\mathbf{C}}_{t|T}\left(I - \hat{\mathbf{A}}_{t|T}e^{-i\omega}\right)^{-1}\hat{\mathbf{B}}_{t|T}\hat{\mathbf{B}}'_{t|T}\left(I - \hat{\mathbf{A}}'_{t|T}e^{i\omega}\right)^{-1}\hat{\mathbf{C}}'_{t|T} \ . \tag{9}$$

The low-frequency relationship between deficits over debt and inflation at time $t$ is computed as[8]

$$\hat{b}_{f,t|T} = \frac{S_{\pi,d,t|T}(0)}{S_{d,t|T}(0)} \ . \tag{10}$$

## 2.4 Estimation results for the low-frequency relationship

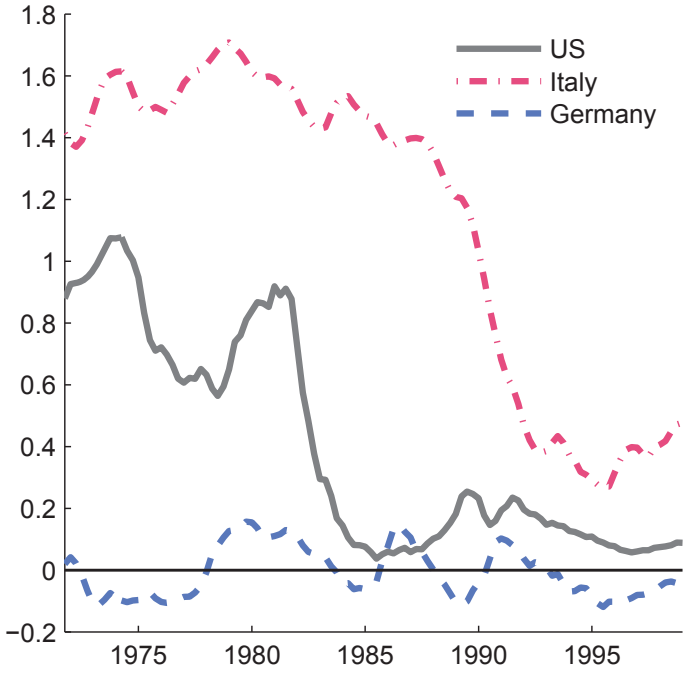
In this section we present our estimation results for the low-frequency relationship between deficits over debt and inflation and relate them to narrative accounts for each country.

### USA

The solid (black) line in Figure 1 presents the evidence for the U.S. The low-frequency relationship between deficits and inflation is high during the 1960s and 1970s, with a prevalent nexus between fiscal financing and monetary expansion. This link falls apart when Paul Volcker enters the scene, establishing an anti-inflationary policy that is backed by the U.S. government. The behavior is well in line with narrative sources, which characterize the interaction between monetary and fiscal policies.

[8]See Sargent and Surico (2011) and Kliem et al. (2015) for a more detailed discussion.

Figure 1: Mean estimates of the low-frequency relationship between inflation and fiscal deficits over debt for Germany, Italy and the U.S. Graphs including probability bands for each country are presented in Appendix B.3.




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, 2014; Bianchi and Melosi, 2013). 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 interaction changes after Paul Volcker became Fed Chairman. As Meltzer (2010) points out, Volcker rebuilt much of the independence and credibility that the Federal Reserve had lost during the two preceding decades. In this regard, Martin (2015) presents the number of meetings at the White House between the U.S. President and the Fed Chairman. He shows that meetings were quite frequent with Presidents Nixon and Ford (1969-1977) and took place four times more often than with the next four presidents put together. Additionally, Martin (2015) shows that President Johnson (1963-1969) met with the Fed Chairman 300 times during his five years in office. Cochrane (2014) argues that, at the same time

Volcker rebuilt the independence of the Federal Reserve, the fiscal authority implemented fiscal reforms to back outstanding government debt with future primary surpluses. Using the narrative account of Romer and Romer (2010), Kliem et al. (2015) also reason that the public expected the fiscal authority to accommodate the actions of the central bank.

**Germany**

The dashed (blue) line in Figure 1 presents the low-frequency relationship between deficits over debt and inflation for Germany. We observe that the low-frequency relationship is around zero over the whole sample. The result that there is no relationship between the variables of interest is also well in line with narrative sources. Sargent (1982), among other authors, describes the key event in shaping German's attitude towards inflation after 1923: the hyperinflation between 1921 and 1923. This event led to the loss of most private savings and high inflation aversion in Germany.[9] Consequently, in order to create trust in the newly issued currency after World War II, the Bundesbank as well as its predecessor the Bank deutscher Länder were strongly committed to maintaining price stability.

As Beyer, Gaspar, Gerberding, and Issing (2013) describe in their thorough study, the Bundesbank adopted a monetary targeting framework in 1974 following the breakdown of the Bretton-Woods system.[10] Consequently, inflation in Germany peaked at 7.8% in the mid-1970s but remained stable at lower rates afterwards. Though the Bundesbank was also involved in buying government bonds during the mid-1970s, the amount bought by the Bundesbank never exceeded 0.2 percent of GDP. Even the second oil price shock did not lead to high inflation rates at the end of the 1970s and early 1980s.

Furthermore, Germans well understood that hyperinflation in the 1920s was caused by the central bank monetizing the government debt accumulated during and after World War I. Thus, the Bundesbank enjoyed independence from the fiscal authority early on. Throughout the time period we consider, there were no strong attempts by the fiscal authority to weaken the Bundesbank's independence. Here, German reunification is a case in point. During reunification the currency of East Germany (GDR) was exchanged above market value into West German currency. The favorable exchange rate for the GDR was politically motivated. If there ever was an opportunity to force the Bundesbank to accommodate the action of the fiscal authority, it would have been this moment of patriotic euphoria. It did not happen. In order to maintain price stability after the increase in the monetary aggregate, the Bundesbank raised interest rates sharply, contributing to the recession in 1993.

[9]This widely accepted view is established in more detail by Issing (2005).

[10]See Benati and Goodhart (2010) and Bordo and Siklos (2015) for further descriptions of the conduct of monetary policy in Germany. Both studies are consistent with the account given in this paragraph.

**Italy**

The point-dashed (red) line in Figure 1 reports the low-frequency link between fiscal stance and inflation for Italy, which is positive and remarkably high during the 1970s and 1980s. The relationship suddenly drops to moderate levels at the end of the 1980s and in the early 1990s. As in the case of Germany and the U.S., the narrative source relates the times of the high low-frequency relationship to a regime of fiscal dominance and the times of the decrease in the low-frequency relationship to an increase in the independence of the central bank associated with an increasingly responsible fiscal authority.

More precisely, the interaction between monetary and fiscal policies in Italy in the 1970s is characterized by fiscal dominance. During this period, Banca d'Italia had to act as a residual buyer at treasury bills auctions and thus monetize the public debt. The resulting high inflation rates in the 1970s and the entry into the European Monetary System (EMS) in 1979 led to a change in the interaction between monetary and fiscal policies: the *divorce* of the Banca and the Tresoro in 1981. More precisely, both institutions agreed that the central bank would gradually become independent.[11] The entry into the EMS was associated with the EMS serving as an inflation-stabilizing device.[12]

Although the *divorce* event marks a significant change in the interaction of these institutions, "one should be careful not to jump to the conclusion that 1981 represented a sharp breaking point between the previous regime of fiscal dominance and the new regime of central bank independence" (Fratianni and Spinelli, 1997). The view of the authors is corroborated by the fact that inflation rates in Italy remained among the highest in the EMS. Furthermore, the gradual independence of the central bank was not supported by the fiscal authority. As Bartoletto, Chiarini, and Marzano (2013) point out, the Italian government continued to run deficits in the early 1980s instead implementing fiscal reforms to back outstanding government debt with future primary surpluses. Only from the mid-1980s did debt stabilization become a target of the fiscal authority (see Balassone, Francese, and Pace, 2013).

The final regime change in Italy occurred in the context of efforts by Italy to join the European Monetary Union (EMU). Banca d' Italia became operationally independent in 1992. Moreover, in order to comply with the Maastricht Treaty, the law which could force the monetary authority to act as a residual buyer at treasury bills auctions was abolished de jure.

[11] Carlo Ciampi, the Governor of the Banca d'Italia at that time expressed this agreement in a speech to shareholders on May 30th, 1981.

[12] Giavazzi and Pagano (1991) provide an economic model for the effects of adopting a fixed exchange rate regime on inflation.

**Comparison and summary**

The results in this section show that the low-frequency relationship between public deficits over debt and inflation varies across countries and times. It is remarkably consistent with the narrative sources in the sense that a period of fiscal dominance and an accommodative central bank is related to a high estimate. By contrast, an independent central bank and a fiscal authority which is concerned with raising sufficient primary surpluses is related to a low estimate. In the next section, we will demonstrate that the evolution of the low-frequency relationship is indeed related to the interaction between monetary and fiscal policies.

# 3 Structural interpretation

In this section, we interpret our results structurally. Our aim is to show that the evolution of the low-frequency relationship is related to a change in the interaction between monetary and fiscal policies. We start by establishing that the change in the low-frequency relationship in Italy as well as in the U.S. is not due to a change in the volatilities of the underlying structural shocks, but due to change in the systematic part of the economy. To do so, we employ the TVP-VAR model and conduct a counterfactual analysis. In a next step, we use a DSGE model to identify all potential changes in the systematic part of the economy, which can explain the actual as well as the counterfactual low-frequency relationship. Due to data limitations, this exercise is restricted to the U.S..

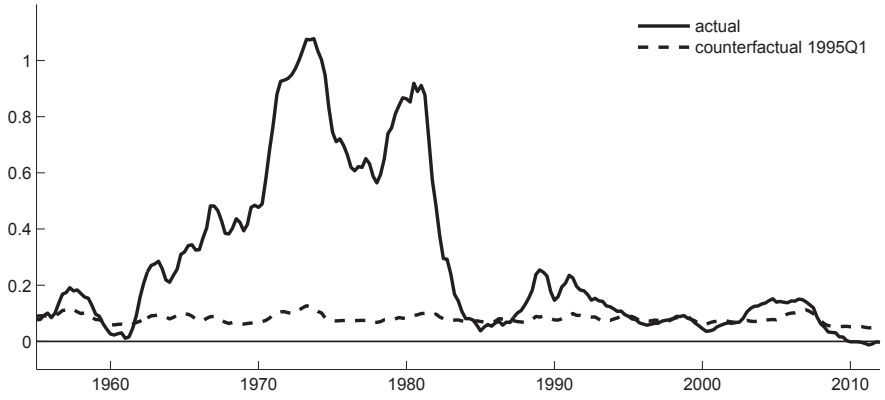
## 3.1 Counterfactual TVP-VAR model analysis

This section tackles the question of whether the change in the low-frequency relationship in Italy and the U.S. is due to a change in the volatilities of the shocks or due to a change in the systematic part of the economy. For the sake of readability, we restate the TVP-VAR model:

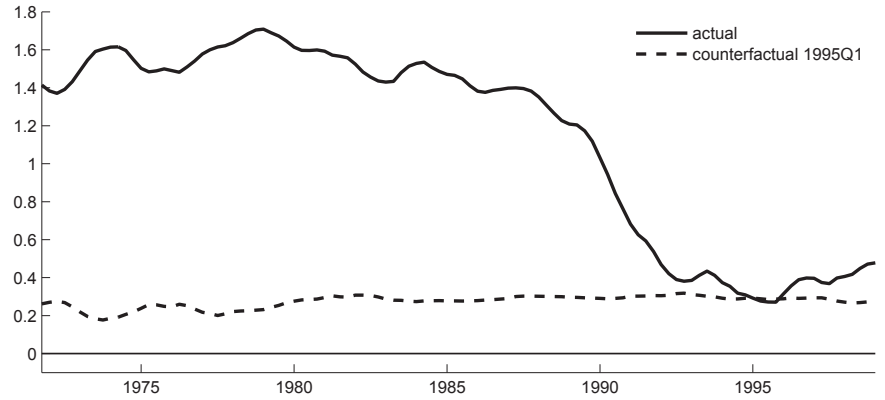
$$y_t = c_t + \sum_{j=1}^{p} A_{j,t} y_{t-j} + B_t \epsilon_t \quad \epsilon_t \sim \mathcal{N}(0, H_t) \tag{11}$$

In this model the coefficient matrices $A_{j,t}$ and $B_t$ represent the systematic part of the economy. The matrix $H_t$ contains the volatilities of the shocks. We start by fixing the systematic behavior of the U.S. as well as the Italian economy to the first quarter of 1995. 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 low-frequency relationship using

Equation (10). The results are displayed in Figure 2. It shows that had the systematic part of the U.S. and the Italian economy been as in 1995Q1, the high low-frequency relationship between inflation and fiscal deficits over debt would have occurred neither in the U.S. nor in Italy.



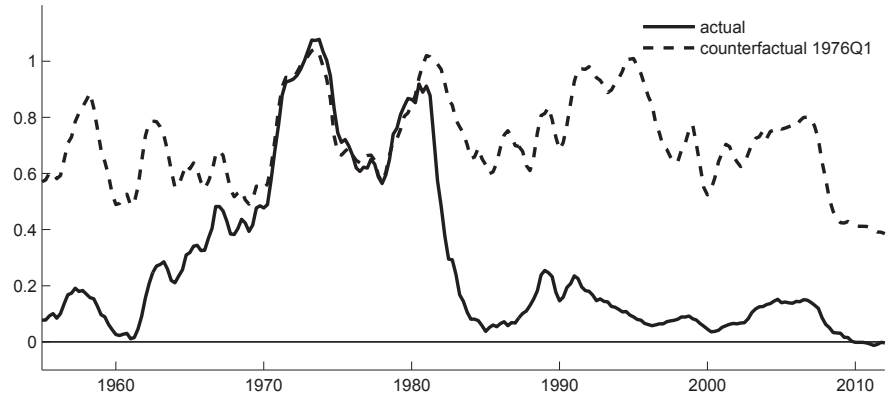

(a) Counterfactual U.S.



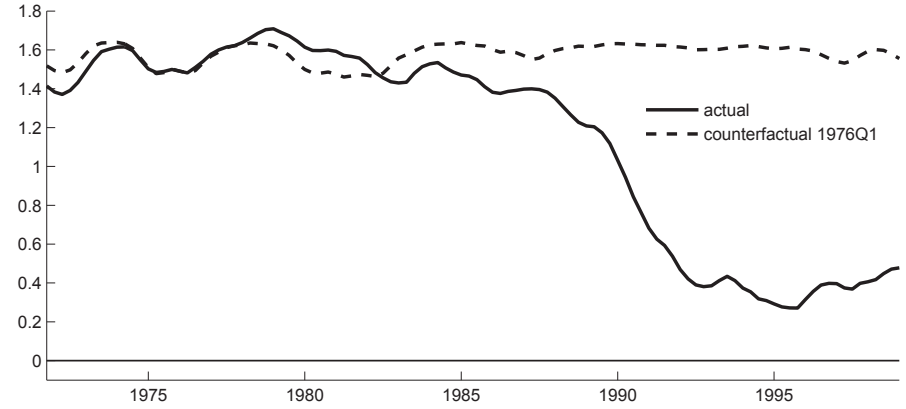

(b) Counterfactual Italy

Figure 2: Counterfactual – the systematic part of the economy is fixed to the first quarter in 1995.

In a second experiment, we fix the systematic part of both economies to the first quarter in 1976. Again, we draw realizations for $\mathbf{A}_{1976.1}$ and $\mathbf{B}_{1976.1}$ out of their posterior distributions, the matrix $(\mathbf{H}_t)$ from its posterior distribution at each point in time and calculate the low-frequency relationship. Figure 3 presents the results. It shows that had the U.S. and Italian economy been in the same state they were in 1976Q1 all the time, the low-frequency relationship after 1980 would not have been around zero in the U.S. or would not have substantially declined in Italy.




(a) Counterfactual U.S.




(b) Counterfactual Italy

Figure 3: Counterfactual – the systematic part of the economy is fixed to the first quarter in 1976.

Finally, we take from this counterfactual exercise that shifts in the low-frequency relationship can be explained more convincingly by changes in the propagation of shocks than by variations in the volatility of shocks. To further interpret the estimation results from the reduced form TVP-VAR model, we employ a DSGE model in the next step.

## 3.2 A DSGE model

In this section, we first set up the DSGE model. Afterwards, we estimate the DSGE model by paying special attention to the low-frequency relationship between deficits over debt and inflation. Finally, we conduct a counterfactual experiment to demonstrate that changes of the interaction between monetary and fiscal policies can explain our empirical findings.

**Model description**

As a starting point, we consider the DSGE model set up by Bhattarai, Lee, and Park (2015). This DSGE model is a New-Keynesian model and exhibits many features which are potentially important for characterizing the low-frequency relationship between fiscal deficits and inflation: trend inflation, partial dynamic indexation in price setting, and external habit formation in consumption. In addition, we follow Bianchi and Ilut (2014) and add long-term government debt. For a thorough and detailed description of the properties of the DSGE model, we refer the reader to these two papers. Below, we will employ standard notation for the variables and parameters of the DSGE model. This notation will partly contradict the one hitherto employed in the paper. We therefore point out that the following notation applies only to the DSGE model and is valid only in Section 3.2 and Section 3.3.

The household $j$ maximizes the utility function:

$$E_0 \left[ \sum_{t=0}^{\infty} \beta^t \left[ \log \left( C_t^j - hC_{t-1} \right) - \frac{\left( H_t^j \right)^{1+\vartheta}}{1+\vartheta} \right] \right] \tag{12}$$

subject to the following budget constraint:

$$P_t C_t^j + P_t^m B_t^m + P_t^s B_t^s = W_t \left( j \right) H_t^j + B_{t-1}^s + \left( 1 + \rho_m P_t^m \right) B_{t-1}^m + P_t D_t - T_t \ , \tag{13}$$

where $C_t^j$ is consumption of household $j$, $C_t$ is aggregate consumption, and $H_t^j$ denotes hours worked. Furthermore, $D_t$ denotes real dividends paid by the firms, $P_t$ the aggregate price index, $W_t \left( j \right)$ the competitive nominal wage, and $T_t$ lump-sum taxes net of government transfers. The parameter $\beta$ denotes the discount factor, $h$ the degree of external habit formation, and $\vartheta$ the inverse of the Frisch elasticity of labor supply. Finally, we follow Eusepi and Preston (2013) and Woodford (2001) and assume that the household has access to two types of government debt. First, it holds one-period government bonds, $B_t^s$, which are assumed to be in zero net supply and has price $P_t^s$. This bond pays a risk-less interest rate $R_t$. Additionally, the household holds a more general portfolio of government bonds, $B_t^m$, in non-zero net supply with price $P_t^m$. This portfolio has the payment structure $\rho_m^{T-(t+1)}$ for $T > t$ and $0 < \rho_m < 1$. The value of such a portfolio issued in period $t$ in any future period $t+k$ is $P_{t+k}^{m-k} = \rho_m^k P_{t+k}^m$. It can be interpreted as a portfolio of infinitely many bonds, with weights along the maturity structure, which is given by $\rho_m^{T-(t+1)}$. Varying the parameter $\rho_m$ varies the average maturity of debt.

The final good $(Y_t)$ is produced by firms assembling intermediate goods $(Y(i))$ using a Dixit-Stiglitz production technology, $Y_t = \left( \int_0^1 Y_t \left( i \right)^{\frac{\varepsilon_t - 1}{\varepsilon_t}} \right)^{\frac{\varepsilon_t}{\varepsilon_t - 1}}$, where $\varepsilon_t$ denotes the time-

varying elasticity of substitution between the intermediate goods with steady state $\bar{\varepsilon}$. The intermediate goods producer $(i)$ has access to the production function:

$$Y_t(i) = A_t H_t(i) \tag{14}$$

where $A$ is an aggregate technological process, which evolves according to the process: $\log\left(\frac{A_t}{A_{t-1}}\right) = \gamma + \hat{z}_t$, $\hat{z}_t = \rho_z \hat{z}_{t-1} + \sigma_z \epsilon_{z,t}$, $\epsilon_{z,t} \sim N(0,1)$.[13]

Price setting in the intermediate goods sector is modeled following Calvo (1983). A firm is allowed to reset its price optimally with probability $1-\theta$ every period. Firms which are not allowed to reset their price, adjust their price according to the partial price indexation rule:

$$P_t(i) = P_{t-1}(i)\, \pi_{t-1}^{\zeta} \bar{\pi}^{1-\zeta} \tag{15}$$

where $\zeta$ measures the degree of price indexation. Firms which are allowed to reset their price, choose a common $P_t^*$ to maximize the following present value of future profits:

$$E_t \sum_{t=0}^{\infty} \theta^k \frac{\partial U_{t+k}}{\partial C_{t+k}^j} \frac{\partial C_t^j}{\partial U_t} \left[ P_t^* X_{t,k} - \frac{W_{t+k}(i)}{A_{t+k}} \right] Y_{t+k}(i), \tag{16}$$

where

$$X_{t,k} = \begin{cases} \Pi_{k=1}^{\infty} \pi_{t+k-1}^{\zeta} \bar{\pi}^{1-\zeta}, \; K \geq 1 \\ 1, \; k = 0 \end{cases}. \tag{17}$$

The government sector consists of a monetary authority and a fiscal authority. Under the assumption that one-period debt is in zero net supply, the flow budget constraint of the fiscal authority is given by

$$P_t^m B_t^m = \left(1 + \rho_m P_t^m\right) B_{t-1}^m - T_t + G_t + S_t \;, \tag{18}$$

where $P_t^m B_t^m$ is the market value of debt. Furthermore, $T_t$ and $G_t$ represent fiscal authority's tax revenues and expenditures, respectively. Following Bianchi and Ilut (2014), we introduce the term $S_t$ which is meant to capture a series of features which are not explicitly modeled (e.g. term premium and maturity structure).[14] Below, we normalize the fiscal authority's

[13] Throughout the paper, all variables indicated by "$\wedge$" represents the log-linear deviation of the detrended variable from its corresponding steady state, $\hat{x} = \log\left((X_t/A_t)/\left(\bar{X}/\bar{A}\right)\right)$. By contrast, all variables normalized by GDP, indicated by "$\sim$", are linearized around its steady state, $\tilde{x}_t = X_t - \bar{X}$.

[14] Moreover, given our vector of observables which includes the market value of debt as well as primary surpluses, this shock is necessary to avoid stochastic singularity of the likelihood function when estimating the model.

flow budget constraint by GDP:

$$b_t^m = \frac{b_{t-1}^m R_{t-1,t}^m}{\pi_t Y_t / Y_{t-1}} - \tau_t + g_t + s_t \tag{19}$$

where $b_t^m = \left(P_t^m B_t^m\right) / \left(P_t Y_t\right)$, $g_t = G_t/Y_t$, $\tau_t = T_t/Y_t$, and $s_t = S_t/Y_t$. Moreover, $R_{t-1,t}^m = \left(1 - \rho_m P_t^m\right) / P_{t-1}^m$ is the realized return of the bond portfolio. We assume that $s$ is an exogenous process with $\tilde{s}_t = \rho_s \tilde{s}_{t-1} + \sigma_s \epsilon_{s,t}$, $\epsilon_{s,t} \sim N(0,1)$. What is more, the fiscal authority sets its two fiscal instruments, government spending and tax revenues, according to the following simple rules:

$$\hat{g}_t = \rho_g \hat{g}_{t-1} + \sigma_g \epsilon_{g,t} \qquad \epsilon_{g,t} \sim N(0,1) \tag{20}$$

$$\hat{\tau}_t = \rho_\tau \hat{\tau}_{t-1} + (1 - \rho_\tau)\, \phi_b b_{t-1}^m + \sigma_\tau \epsilon_{\tau,t} \qquad \epsilon_{\tau,t} \sim N(0,1) \tag{21}$$

Finally, the monetary authority sets the nominal interest rate according to the following log-linearized rule:

$$\hat{r}_t = \rho_r \hat{r}_{t-1} + (1 - \rho_r) \left[\phi_\pi \hat{\pi} + \phi_Y \left(\hat{y}_t - \hat{y}_t^N\right)\right] + \sigma_R \epsilon_{R,t} \quad \epsilon_{R,t} \sim N(0,1)\ . \tag{22}$$

This rule features interest-rate smoothing and a systematic response to deviations of inflation from its steady state and to deviations of output from its natural level $\hat{y}_t^N$.

**Estimation of the DSGE model**

We estimate the DSGE model using U.S. data from 1982Q4 to 2008Q2. We choose this episode because it is widely recognized as a time span when fiscal policy ensured the stability of real debt by adjusting future primary surpluses and monetary policy followed the Taylor principle to stabilize inflation.

To estimate the parameters of the model, we use six quarterly time series as observables. As in the TVP-VAR model, we use primary deficits over lagged debt and annual inflation. Moreover, we employ per capita output growth, the annualized federal funds rate, the debt-to-output ratio, as well as the government spending-to-output ratio. Following Bhattarai et al. (2015), we use the following definition for our variables. Per capita output is the sum of personal consumption of nondurables & services and government consumption divided by civilian non-institutional population. For government spending, we use the time series government purchases. The annualized federal funds rate is calculated as quarterly averages of the daily effective fed funds rate. Annual inflation is calculated as the year-to-year change in the log-GDP deflator. Primary deficits are calculated as government purchases minus tax revenues. The last-named variable is defined as the sum of current tax receipts and

contributions for government social insurance. For government debt, we use the market value of privately held gross federal debt.[15]

By using the aforementioned definitions for primary surpluses and government debt, we slightly deviate from the definitions used in the first part of the paper. This stems from the need to adjust the observable variables to their counterparts in the DSGE model. Importantly, changing the definitions of primary deficits and government debt in this manner does not change the evolution of the low-frequency relationship between primary deficits over lagged debt and inflation (see Kliem et al., 2015). Finally, we remove the linear trend from primary deficits over debt, while keeping the remaining variables unchanged. An overview of the measurement equations can be found in Appendix C.

We calibrate the parameters which are not identified given the data set. In particular, we calibrate the Frisch elasticity of labor supply to $1/\vartheta = 0.6$, the steady state of the elasticity of substitution between the intermediate goods $\bar{\epsilon} = 6$, and the average maturity of government debt to 5 years (which is controlled by the parameter $\rho_m$).

Since the focus of the paper is on the low-frequency relationship between inflation and fiscal deficits over debt, we pay special attention to this measure when estimating the DSGE model. To do so, we use an "endogenous prior" approach similar to Del Negro and Schorfheide (2008).[16] First, we specify a set of initial priors, $p(\omega)$, where the priors are independent across parameters. Second, we use a pre-sample to calculate low-frequency characteristics of the variables of interest, which are the point estimates for the spectrum and cross-spectrum of primary deficits over debt and annual inflation at frequency 0.[17] We collect the statistics in a vector $\hat{S}$. For any parameter vector of the DSGE model, $\omega$, we calculate the same statistics. Therefore, we denote $S_{DSGE}(\omega)$ as a function of $\omega$ and impose the following relationship:

$$\hat{S} = S_{DSGE}(\omega) + \eta, \tag{23}$$

where $\eta$ is a vector of measurement errors. Following Del Negro and Schorfheide (2008), we derive a quasi-likelihood function $\mathcal{L}\left(S_M(\omega)|\hat{S}\right) = p\left(\hat{S}|S_M(\omega)\right)$[18] and combine this function with the set of initial prior distributions $p(\omega)$ to obtain the a conditional distribution which

[15] All time series are publicly available from the FRED II database of the Federal Reserve Bank of St. Louis, the Federal Reserve Bank of Dallas, or from the National Income and Product Accounts (NIPA).

[16] Other papers in the literature which use different but related approaches are, for example, Christiano, Trabandt, and Walentin (2011) and Kliem and Uhlig (2013).

[17] In practice, we follow Christiano et al. (2011) and use the actual sample as our pre-sample as no other suitable data is available.

[18] In our application, the function $\mathcal{L}\left(S_M(\omega)|\hat{S}\right)$ is constructed under the assumption that the error terms are independently and normally distributed.

reflects our "endogenous priors":

$$p\left(\omega|\hat{S}\right) \propto p\left(\hat{S}|S_{DSGE}\left(\omega\right)\right) p(\omega) \tag{24}$$

While the initial priors are independent across parameters, the "endogenous priors" for the actual sample are not independent across parameters. We combine this endogenous prior distribution with the likelihood in order to obtain the posterior distribution. The elements of $\hat{S}$ are taken from a VAR model. Instead of computing the median estimate of the TVP-VAR model from 1984-2009, we estimate a time-invariant VAR model with six lags. This has the advantage that the uncertainty around the estimate is smaller.[19] Table 1 shows the prior distributions for the elements ($\tilde{S}$). The two parameters can be interpreted as $\hat{S}$ value and the standard deviation of $\eta$. The estimation results for the parameters of the DSGE model are also reported in Table 1. Additionally, the table shows the posterior predictions of the low-frequency relationship between fiscal deficits over debt and inflation, $\beta_{\pi d}$. These values are within the probability bands of the posterior distribution of the TVP-VAR model. Hence, the estimated DSGE model performs well in predicting the low-frequency characteristics of interest for the U.S. during the Great Moderation.

## 3.3 Counterfactual DSGE model analysis

In this section, we investigate which feature of the systematic part of the economy is the main driving force behind the decrease in the low-frequency relationship between fiscal deficits over debt and inflation in the U.S. In our exercise, we aim to replicate the counterfactual TVP-VAR model exercise in Section 3.1. The estimation results in Section 3.2 show that the DSGE model is able to replicate the mean of the low-frequency relationship between 1984 and 2009. The counterfactual analysis in Section 3.1 implies that the change in the low-frequency relationship is not due to a change in the volatilities of the structural shocks. Consequently, we now fix the standard deviations of the shocks in the DSGE model to their posterior means. For the remaining structural parameters we conduct a prior predictive analysis with respect to the low-frequency relationship between fiscal deficits over debt and inflation. The prior predictive analysis reports the probability distribution of the low-frequency relationship between fiscal deficits over debt and inflation that a specific model can produce before it is confronted with any data.

As in Leeper, Traum, and Walker (2015), we employ our prior predictive analysis for two different regimes of monetary and fiscal interaction. In particular, we will concentrate on

[19]The results of this estimation are consistent with the average of the corresponding estimates based on our five-variable TVP-VAR(2) model.

| | Posterior distribution | | | Prior distribution | | |
|---|---|---|---|---|---|---|
| | mean | 5% | 95% | density | mean | std |
| $\rho_z$ | 0.9562 | 0.9463 | 0.9668 | Beta | 0.80 | 0.10 |
| $\rho_\mu$ | 0.4602 | 0.3479 | 0.5676 | Beta | 0.80 | 0.10 |
| $\rho_g$ | 0.9766 | 0.9661 | 0.9878 | Beta | 0.80 | 0.10 |
| $\rho_s$ | 0.7605 | 0.7394 | 0.7834 | Beta | 0.80 | 0.10 |
| $\rho_r$ | 0.8863 | 0.8585 | 0.9143 | Beta | 0.80 | 0.10 |
| $\phi_\pi$ | 2.1121 | 1.7061 | 2.4951 | Gamma | 2.00 | 0.25 |
| $\phi_y$ | 0.1382 | 0.0665 | 0.2085 | Gamma | 0.12 | 0.05 |
| $\rho_\tau$ | 0.8978 | 0.8479 | 0.9528 | Beta | 0.80 | 0.10 |
| $\phi_b$ | 0.0746 | 0.0441 | 0.1049 | Gamma | 0.07 | 0.02 |
| $\zeta$ | 0.1822 | 0.0698 | 0.2773 | Beta | 0.30 | 0.10 |
| $\theta$ | 0.8437 | 0.7840 | 0.9053 | Beta | 0.50 | 0.10 |
| $h$ | 0.6849 | 0.6021 | 0.7662 | Beta | 0.50 | 0.10 |
| $\bar{g} \cdot 100$ | 21.0759 | 20.5251 | 21.6281 | Normal | 21.00 | 2.00 |
| $\bar{b}^m \cdot 100/4$ | 45.6909 | 42.7379 | 48.9131 | Normal | 48.00 | 2.00 |
| $(\bar{\pi} - 1) \cdot 100$ | 0.5997 | 0.5264 | 0.6641 | Normal | 0.64 | 0.10 |
| $(\Gamma - 1) \cdot 100$ | 0.4854 | 0.3481 | 0.6340 | Normal | 0.57 | 0.10 |
| $(1/\beta - 1) \cdot 100$ | 0.1920 | 0.1321 | 0.2459 | Gamma | 0.25 | 0.05 |
| $\sigma_z \cdot 100$ | 0.1846 | 0.1494 | 0.2193 | InvGamma | 0.50 | Inf |
| $\sigma_\mu \cdot 100$ | 0.1003 | 0.0829 | 0.1165 | InvGamma | 0.50 | Inf |
| $\sigma_r \cdot 100$ | 0.1364 | 0.1193 | 0.1522 | InvGamma | 0.50 | Inf |
| $\sigma_\tau \cdot 100$ | 0.4588 | 0.3981 | 0.5202 | InvGamma | 0.50 | Inf |
| $\sigma_g \cdot 100$ | 0.1570 | 0.1386 | 0.1755 | InvGamma | 0.50 | Inf |
| $\sigma_s \cdot 100$ | 1.4479 | 1.2771 | 1.5996 | InvGamma | 0.50 | Inf |
| $\tilde{S}_d(0)$ | 37.6003 | 35.6516 | 39.5221 | Normal | 37.00 | 3.00 |
| $\tilde{S}_{d\pi}(0)$ | 5.0353 | 4.0048 | 6.1056 | Normal | 7.00 | 1.00 |
| $\tilde{S}_\pi(0)$ | 3.8200 | 2.9245 | 4.80688 | Normal | 2.00 | 1.00 |
| $\beta_{\pi d}$ | 0.1345 | 0.1060 | 0.1626 | - | - | - |

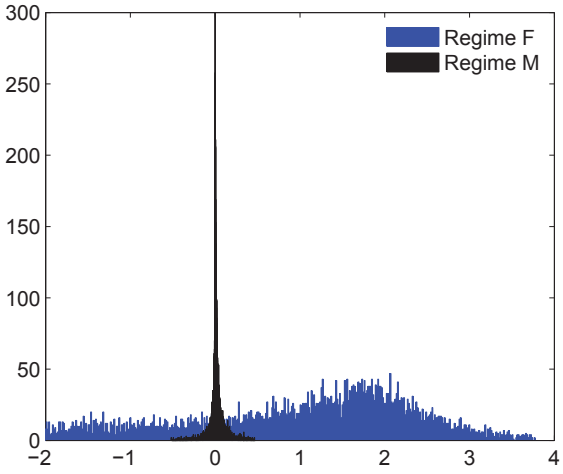
Table 1: Prior and posterior statistics.

uniquely determined bounded rational expectation equilibria. These regimes exhibit either an active monetary authority coupled with an passive fiscal authority (regime $M$) or a passive monetary authority coupled with an active fiscal authority (regime $F$). This terminology of active and passive authorities follows Leeper (1991). While an active fiscal policy is defined as a policy where decisions are not constrained by current budgetary conditions, a passive fiscal policy has to ensure the sustainability of real government debt by raising sufficient primary surpluses. Similarly, a passive monetary authority is constrained by the actions of the active fiscal authority. If the fiscal authority does not raise sufficient primary surpluses to stabilize outstanding government debt, the monetary authority has to set nominal interest rates to maintain the value of government debt. By contrast, under active monetary policy, the monetary authority is free to target inflation by aggressively adjusting nominal interest

rates.

To reflect these two policy regimes, we choose two sets of prior distributions for the policy parameters $\phi_b$ and $\phi_\pi$. While, for regime $M$, we use the same prior distribution as in the estimation (see Table 1) of the DSGE model, we employ the following prior distributions $\phi_b \sim N(-0.025, 0.001)$ and $\phi_\pi \sim N(0.5, 0.1)$ for regime $F$. The prior distributions for the remaining parameters are the same across the regimes and correspond to the distribution specified in Table 1. Figure 4 shows histograms of the predicted low-frequency relationship between fiscal deficits over debt and inflation under both regimes based on 10,000 draws from the corresponding prior distribution. Regime $M$ predicts a quite tight probability distribution for the low-frequency relationship which centered around 0. In contrast, regime $F$ supports a wider range of values for the low-frequency relationships between fiscal deficits over debt and inflation.

Figure 4: Prior predictive analysis of the low-frequency relationship between fiscal deficits over debt and inflation under different regimes. Standard deviation of shocks fixed to their posterior mean.

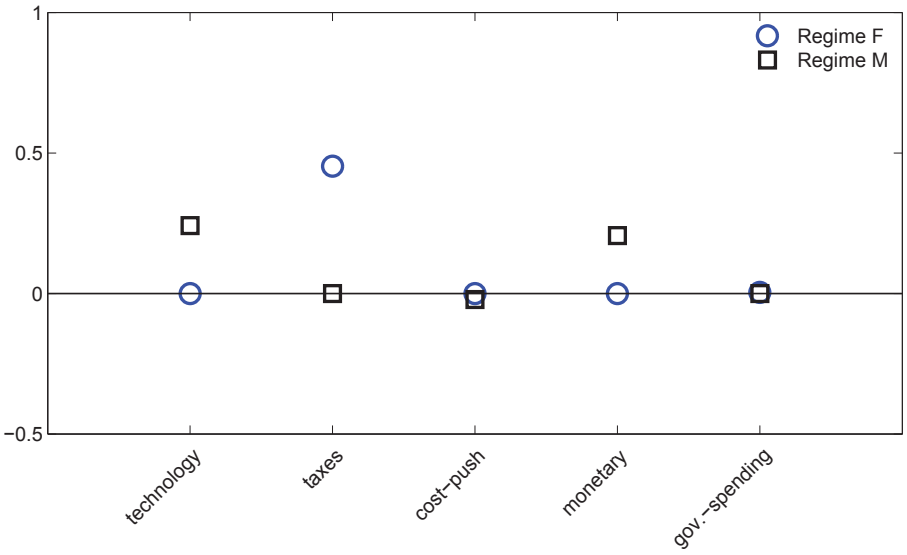



We interpret the results of the prior predictive analysis in the following way. Given the estimated standard deviations, it is very unlikely for regime $M$ to produce high values for the low-frequency relationship between inflation and fiscal deficits over debt. That is, it is very unlikely that the policy regime in 1976Q1 was regime $M$. In contrast, it is very likely for regime $F$ to produce the counterfactual low-frequency relationship from Section 3.1. Hence, the prior predictive analysis pinpoints the fact that the policy parameters, $\phi_b$ and $\phi_\pi$, are the crucial elements of the model in determining the low-frequency relationship between fiscal deficits over debt and inflation. Moreover, this finding is consistent with our narrative evidence about the changes of monetary and fiscal policy interaction. Therefore,

the result can well be applied to Germany and Italy. It implies that Germany was in regime $M$ throughout, while Italy was in regime $F$ until the early 1990s.

In order to uncover the mechanism behind the change in the low-frequency relationship, we investigate how exactly the different policy regimes affect the low-frequency relationship . To that end, we decompose the low-frequency relationship into the contribution of the underlying structural shocks.[20] Figure 5 reports the results as a percentage of the unconditional low-frequency measure. The figure shows that the structural shocks in each regime influence the low-frequency relationship very differently. In regime $M$, the low-frequency relationship is mostly determined by the technology shock and monetary policy. Fiscal policy shocks do not seem to play an important role. In contrast, the low-frequency relationship in regime $F$ is mostly determined by fiscal policy shocks, which (usually) affect the economy differently than the technology shock. Thus, the change in the policy regime matters for the propagation of shocks in the economy, which is most pronounced at lower frequencies.

Figure 5: Percentage contribution of each shock to the low-frequency relationship between fiscal deficits over debt and inflation. The figure shows the median prediction from prior distribution under different regimes. Standard deviation of shocks fixed to their posterior mean.




[20]The unconditional low-frequency measure can be written as the sum of weighted conditional low-frequency measures (see, for example, Kliem et al., 2015; Gambetti and Galí, 2009).

# 4 Conclusion

What determines the relationship between inflation and the fiscal stance at lower frequencies? We have shown that it is the interaction between monetary and fiscal policies. As a starting point, we have contrasted the evolution of the low-frequency relationship between our measure of the fiscal stance, fiscal deficits over debt, and inflation in the U.S. with those of Germany and Italy. For our sample, which ranges from 1965 to 1999, this comparison reveals that the low-frequency relationship is around zero for periods to which narrative accounts assign an independent central bank and a responsible fiscal authority. These characteristics apply to the U.S. after Paul Volcker became chairman of the Federal Reserve, to Italy after it joined the Economic and Monetary Union (EMU) in 1990 and to Germany throughout our sample. In contrast, the low-frequency relationship is high whenever the narrative accounts points to a fiscal authority which did not stabilize its outstanding government debt together with a central bank that accommodated this behavior. These characteristics apply to the 1960s and 1970s in the U.S. and to Italy from the start of the sample up to the early 1990s.

We interpret the estimation results by means of an counterfactual analysis and a DSGE model. We first establish that the change in the low-frequency relationship in Italy and the U.S. is due to a change in the systematic part of the economy and not due to a change in the volatilities of the underlying structural shocks. Afterwards, we employ the DSGE model to show that a change in the interaction between monetary and fiscal policies can well account for the estimated low-frequency relationship after 1984 in the U.S. as well as for the counterfactual low-frequency relationship which would have occurred if the economy had been in same state as in the 1970s. The DSGE model further allows us to uncover how the different policy regimes affect the low-frequency relationship between fiscal deficits over debt and inflation. We find that, in differing policy regimes, different structural shocks determine the low-frequency relationship, i.e., the policy regime matters for the propagation of the structural shocks and thus affects the low-frequency relationship between fiscal deficits and inflation.

Our findings suggests that a structural decomposition of the low-frequency relationship between fiscal deficits over debt and inflation might be helpful in discriminating empirically between different policy regimes. Since the interaction between monetary and fiscal policies determines the low-frequency relationship between fiscal deficits over debt and inflation, the results of this paper corroborate economic models that allow fiscal policy to play a major role in determining of the price level.

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# A Appendix

# B TVP-VAR

## B.1 Prior specification

In this section, we describe our prior choice for the initial conditions of the VAR coefficients and the variance-covariance matrix of the disturbances in the law of motion of the time-varying parameters.

For the priors on the initial conditions of the time-varying VAR coefficients we use multivariate normal distributions, which are parameterized with the OLS estimates obtained from the training sample.

$$\mathbf{a}_0 \sim N\left(\hat{a}^{OLS}, Var\left(\hat{a}^{OLS}\right)\right)$$

Similarly, the prior for the starting values of the off-diagonal elements $B_t$ is

$$\mathbf{b}_0 \sim N\left(\hat{b}^{OLS}, k_b \cdot V\left(\hat{b}^{OLS}\right)\right),$$

where $\hat{b}^{OLS}$ are the off-diagonal element of the OLS estimate of the VAR variance-covariance matrix, $\hat{\Omega}^{OLS}$. $V\left(\hat{b}^{OLS}\right)$ is assumed to be a diagonal with elements equal to the absolute value of the corresponding $\hat{b}^{OLS}$. The hyperparameter $k_b$ is set to 11 which is equal to $(1 + dim\left(\hat{b}^{OLS}\right))$. The prior for the diagonal elements of the VAR variance-covariance matrix is

$$\log h_0 \sim N\left(\log \hat{h}^{OLS}, I_n\right),$$

where $\hat{h}^{OLS}$ are the diagonal elements of $\hat{\Omega}^{OLS}$.

For priors on the variance-covariance matrices of the error terms in the time-varying parameter equations, we use an inverse Wishart distribution , $\mathbf{Q}$ and $\mathbf{S}$. We follow here Cogley and Sargent (2005) and choose the prior for $\mathbf{Q}$ as

$$\mathbf{Q} \sim IW\left(k_Q \cdot Var\left(\hat{a}^{OLS}\right), T\right),$$

where $k_Q = 3.5 \cdot 10^{-4}$ and the degrees of freedom $T = 60$. While the minimum degrees of freedom are equal to $1 + dim\left(\hat{a}^{OLS}\right)$ our choice is slightly higher following Primiceri (2005). Similarly, we follow Primiceri (2005) by specifying the prior for $\mathbf{S}$. Because $\mathbf{S}$ is block diagonal, we choose for each block $i$ the following inverse-gamma distribution

$$\mathbf{S}_i \sim IW\left(k_S \cdot (i+1)\, V\left(\hat{b}_i^{OLS}\right), (i+1)\right),$$

with $k_S = 0.01$ and $V\left(\hat{b}_i^{OLS}\right)$ being diagonal with elements equal to the absolute values of the corresponding blocks of $\hat{b}^{OLS}$. Finally, for the prior for $\sigma_i^2$ which are the diagonal elements of $\mathbf{W}$, we use the following inverse-gamma distribution:

$$\sigma_i^2 \sim IG\left(\frac{10^{-3}}{2},\ \frac{1}{2}\right)$$

## B.2 Convergence statistics of the TVP-VARs

To check the convergence of our sampler for three countries, we have used visual inspections as convergence diagnostics. The visual inspections illustrate how the parameters move through the parameter space, thereby allowing us to check whether the chain gets stuck in certain areas. To visualize the evolution of our parameters, we use running mean plots and trace plots. For lack of space, we present only running mean plots and trace plots for the trace of the variance covariance matrices $\mathbf{Q}$, $\mathbf{W}$ and $\mathbf{S}$. As can be seen in Figures 6 to 11, running mean plots and trace plots both show that the mean of the parameter values stabilize as the number of iterations increases and that the chains of the different TVP-VARs are mixing quite well.

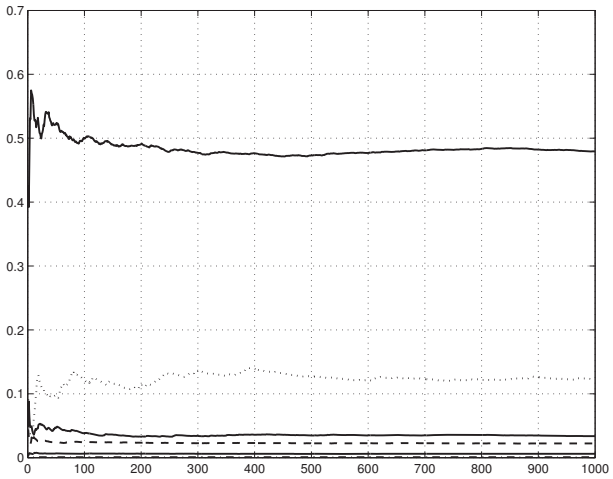



Figure 6: Running Mean Plot for Germany.

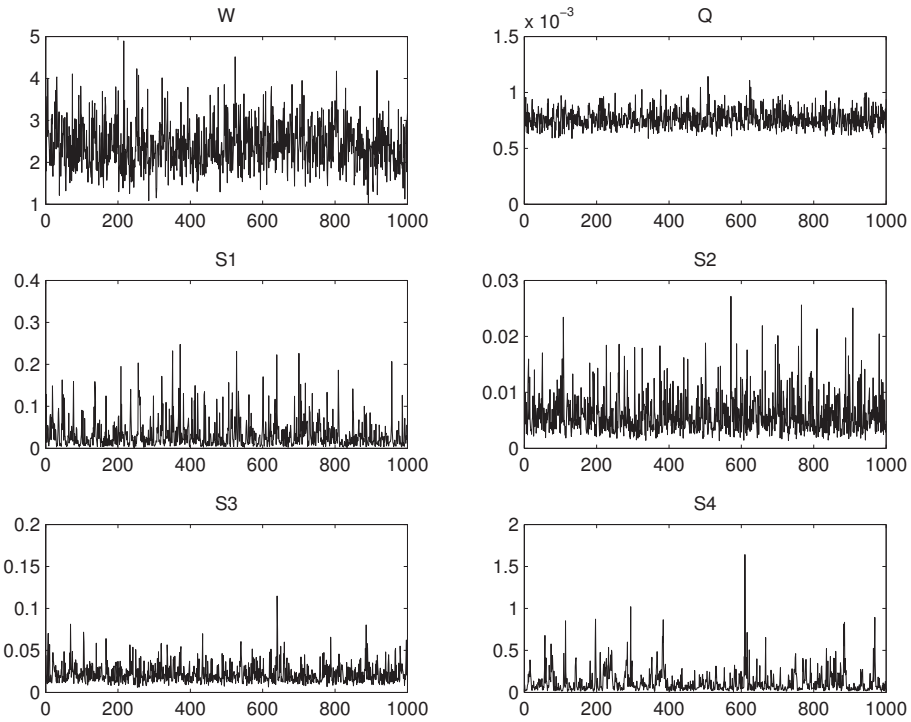



Figure 7: Trace Plot for Germany.

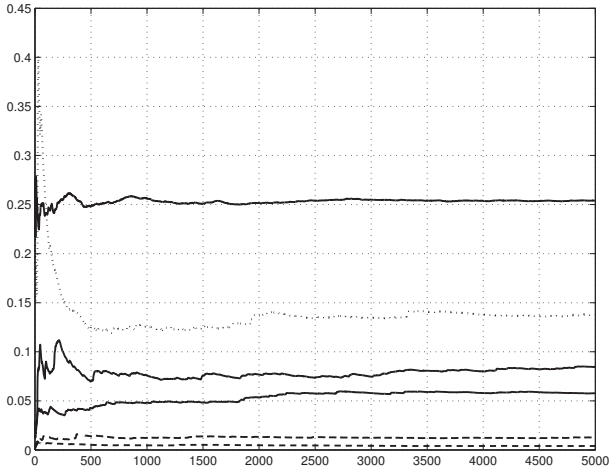


Figure 8: Running Mean Plot for Italy.

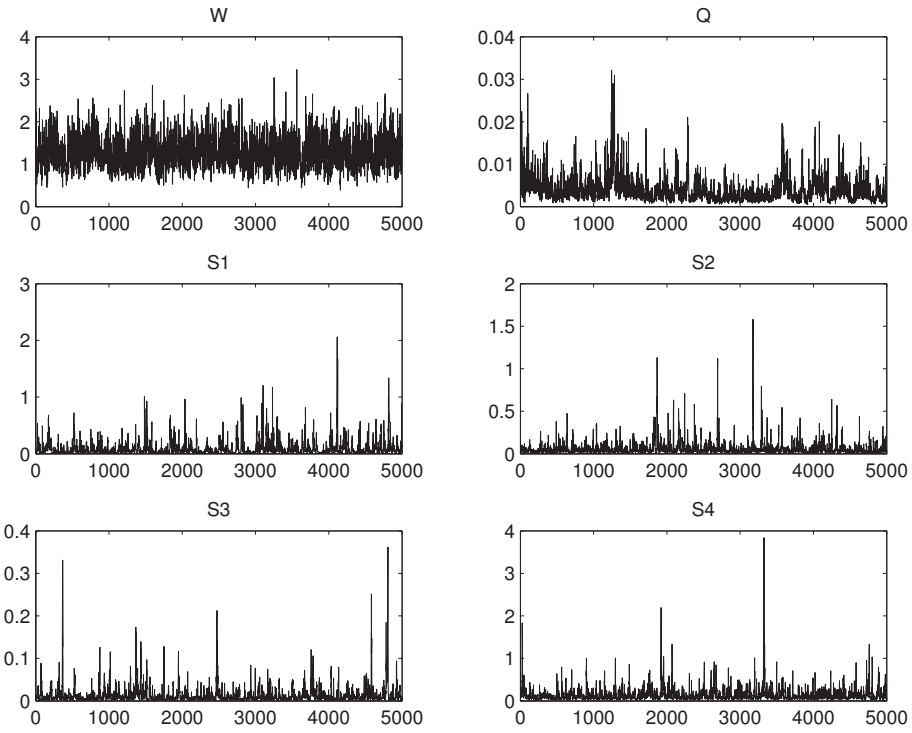



Figure 9: Trace Plot for Italy.

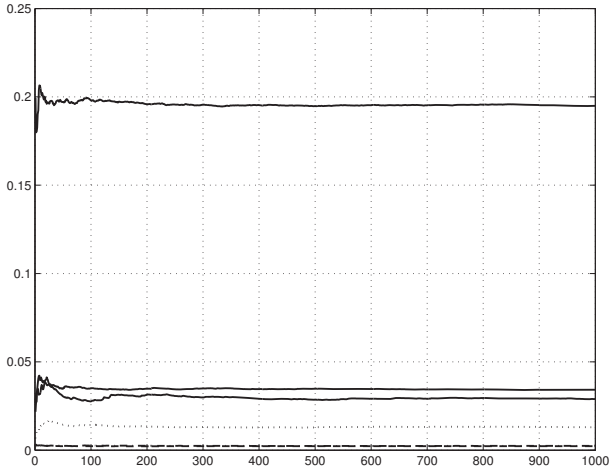


Figure 10: Running Mean Plot for the U.S.

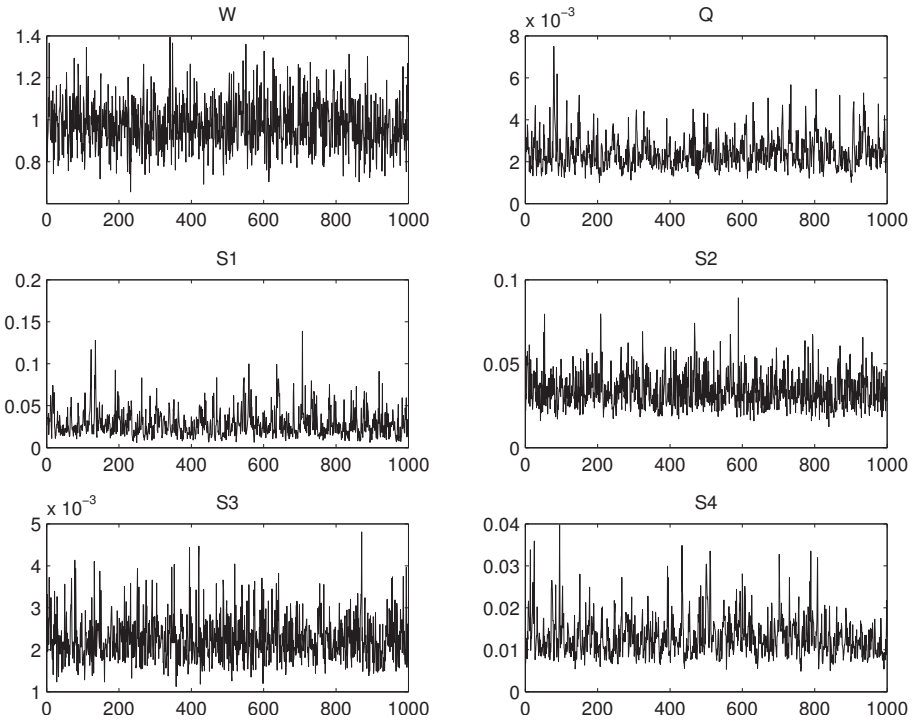
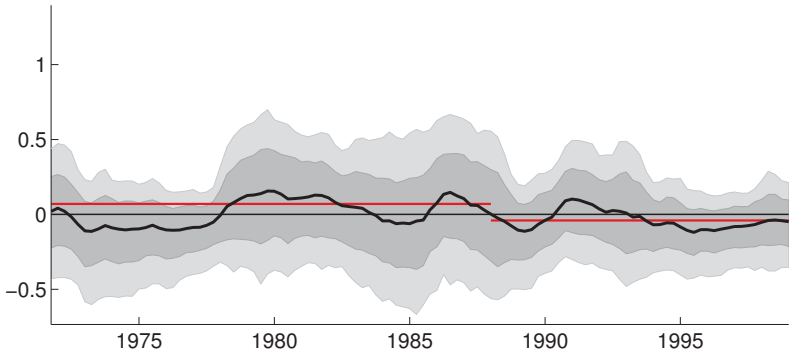
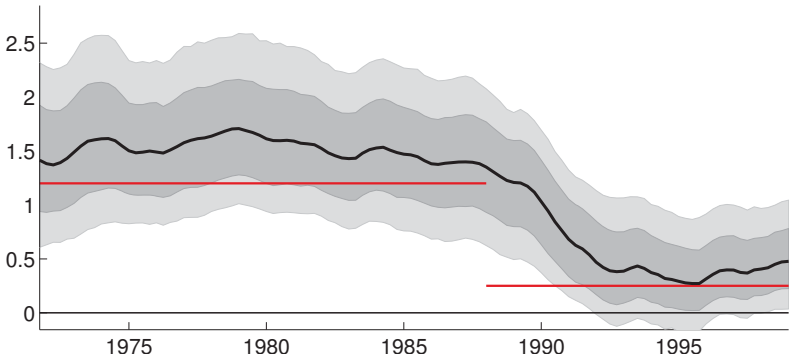



Figure 11: Trace Plot for the U.S.

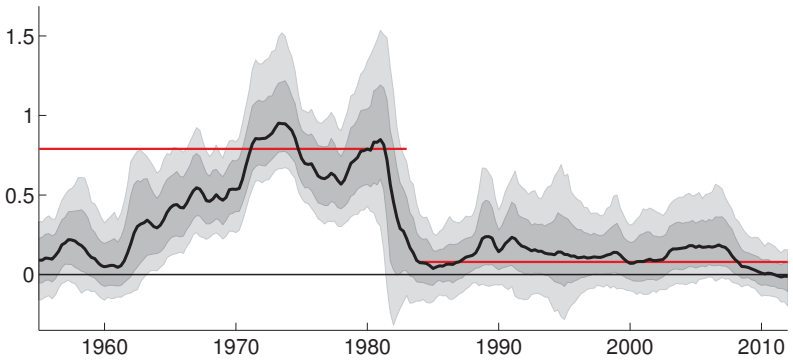
## B.3 Estimation Results




(a) $\hat{b}_f$ for Germany



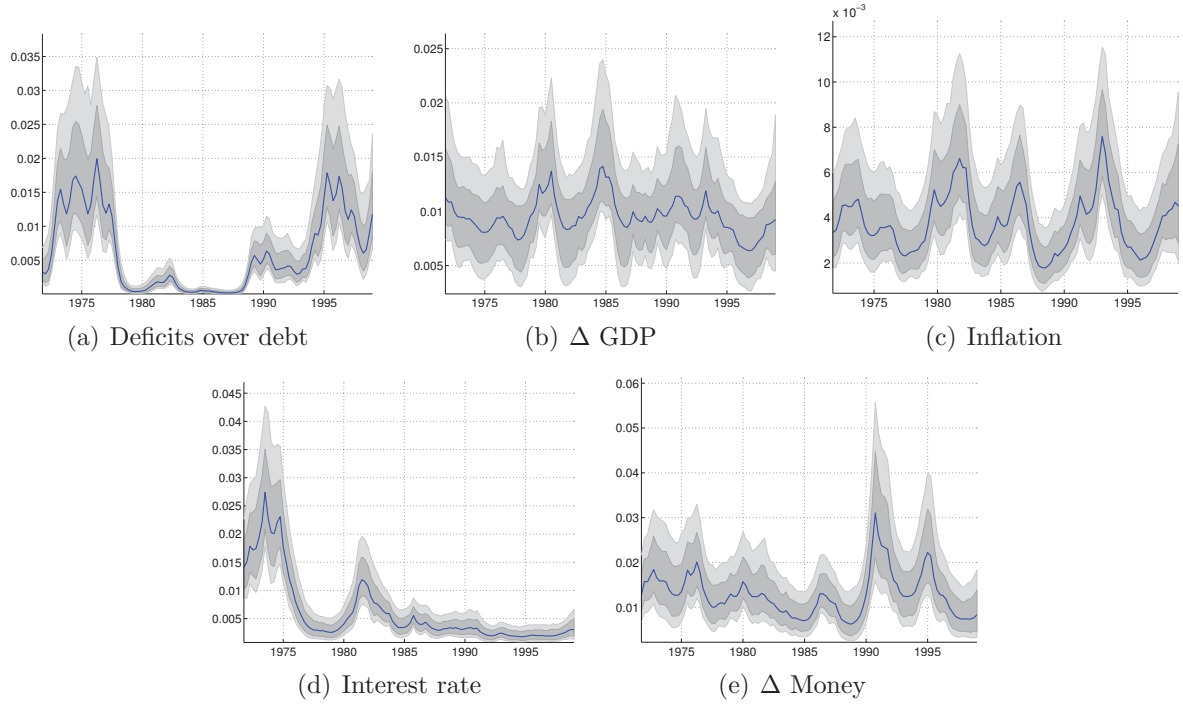

(b) $\hat{b}_f$ for Italy




(c) $\hat{b}_f$ for the U.S.

Figure 12: The black line represents the median, the dark shaded area indicates the 16th and 84th percentiles, and the light shaded area shows the 5th and 95th percentiles of the posterior probability mass of the time-varying regression coefficient of inflation on deficits over debt. Red lines depict slopes of the scatter plots which are based on the OLS regression coefficient of the filtered data in Section D.

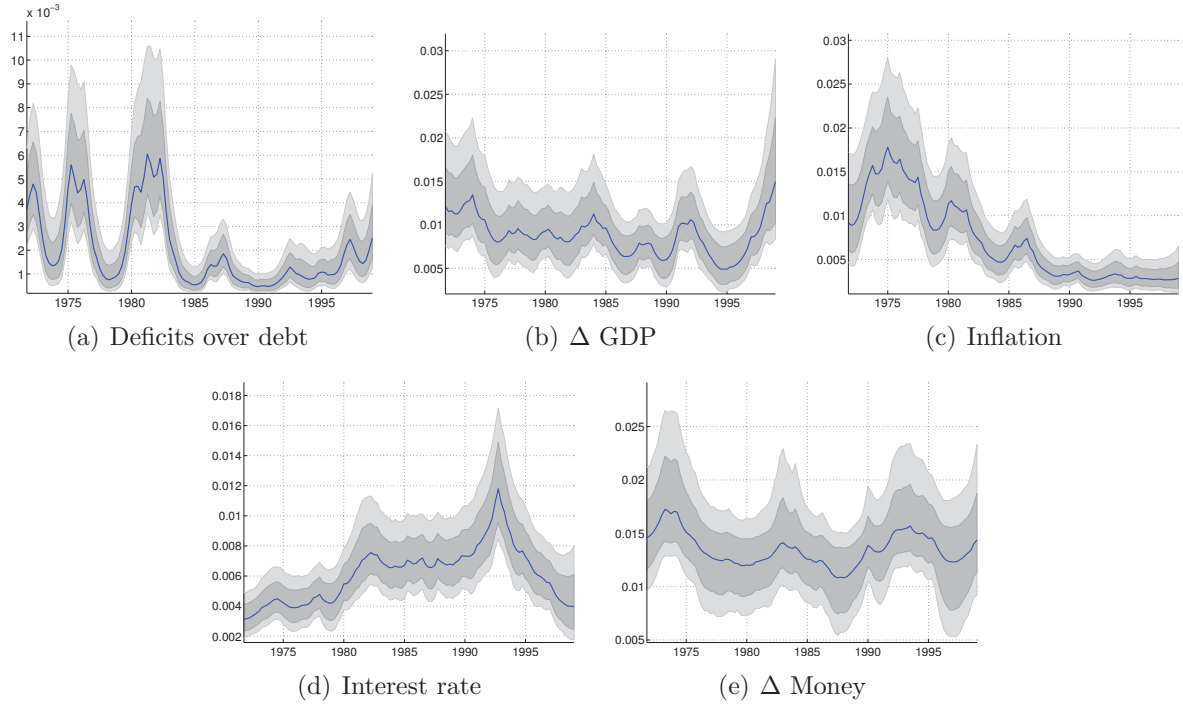
## B.4 Stochastic volatilities




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

(d) Interest rate (e) Δ Money

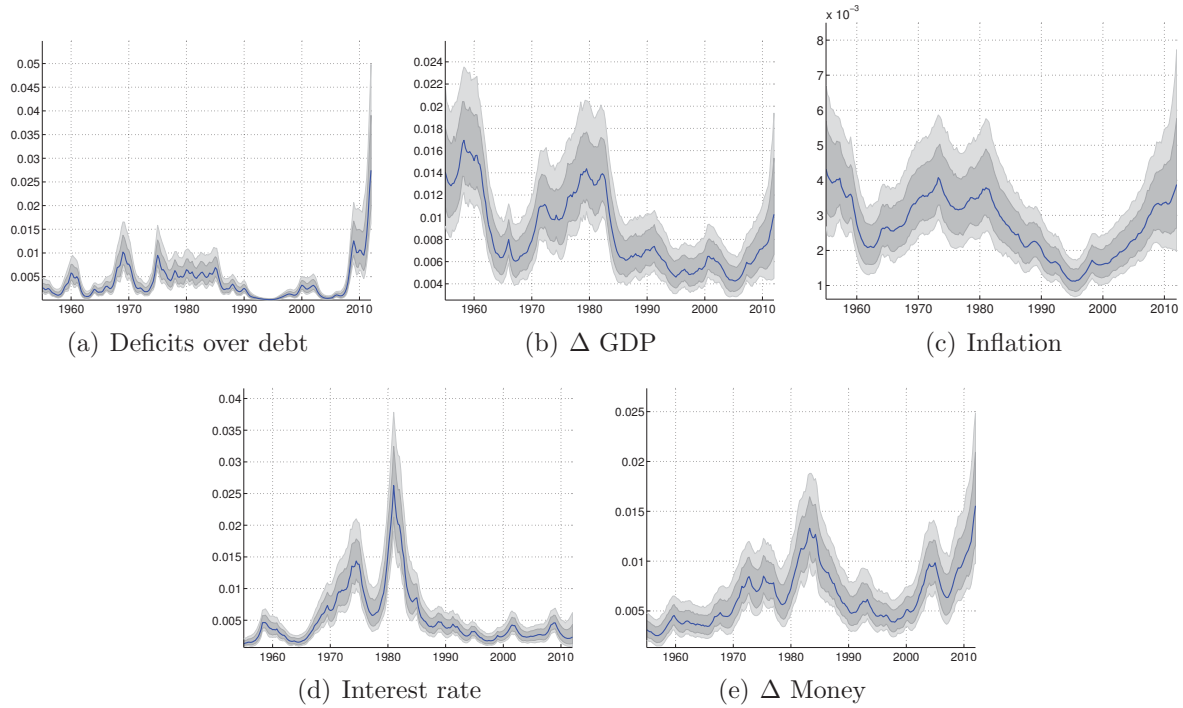
Figure 13: Square roots of stochastic volatility for Germany.




(a) Deficits over debt (b) $\Delta$ GDP (c) Inflation

(d) Interest rate (e) $\Delta$ Money

Figure 14: Square roots of stochastic volatility for Italy.




(a) Deficits over debt (b) $\Delta$ GDP (c) Inflation

(d) Interest rate (e) $\Delta$ Money

Figure 15: Square roots of stochastic volatility for the U.S.

# C DSGE Model

In the following subsection we list the system of equations. All variables are (log- )linearized and detrended if necessary.

consumption Euler equation:

$$\hat{c}_t = \frac{\Gamma}{\Gamma+h}\left(\hat{c}_{t+1} + \rho_z \hat{z}_t\right) + \frac{h}{\Gamma+h}\left(\hat{c}_{t-1} - \hat{z}_t\right) - \frac{\Gamma-h}{\Gamma+h}\left(\hat{r}_t - \hat{\pi}_{t+1}\right), \tag{25}$$

with $\Gamma = \gamma + 1$.

Phillips curve:

$$\begin{aligned}\hat{\pi}_t =& \frac{\zeta}{1+\zeta\beta}\hat{\pi}_{t-1} + \frac{\beta}{1+\zeta\beta}\hat{\pi}_{t+1} \\ &+ \frac{(1-\beta\theta)(1-\theta)}{\theta(1+\zeta\beta)(1+\bar{\varepsilon}\vartheta)}\left(\left(\vartheta + \frac{\Gamma}{\Gamma-h}\right)\left(\hat{y}_t - \hat{y}_t^N\right) - \frac{h}{\Gamma-h}\left(\hat{y}_{t-1} - \hat{y}_{t-1}^N\right)\right) + \hat{\mu}_t,\end{aligned} \tag{26}$$

with $\hat{\mu}_t = -\frac{(1-\beta\theta)(1-\theta)}{\theta(1+\zeta\beta)(1+\bar{\varepsilon}\vartheta)} \cdot \frac{1}{\bar{\varepsilon}-1}\hat{\varepsilon}_t$ which can be interpreted as cost-push shock.

aggregate output:

$$\hat{y}_t = \hat{c}_t + \frac{1}{1-\bar{g}}\tilde{g}_t \tag{27}$$

potential output:

$$\begin{aligned}\hat{y}_t^N =& \frac{h}{\vartheta(\Gamma-h)+\Gamma}\hat{y}_{t-1}^N + \frac{\Gamma}{\vartheta(\Gamma-h)+\Gamma}\frac{1}{1-\bar{g}}\tilde{g}_t \\ &+ \frac{h}{\vartheta(\Gamma-h)+\Gamma}\frac{1}{1-\bar{g}}\tilde{g}_{t-1} - \frac{h}{\vartheta(\Gamma-h)+\Gamma}\hat{z}_t\end{aligned} \tag{28}$$

government budget constraint:

$$\tilde{b}_t^m = \frac{1}{\beta}\tilde{b}_{t-1}^m + \frac{\bar{b}^m}{\beta}\left(\hat{r}_{t-1,t}^m + \hat{y}_{t-1} - \hat{y}_t - \hat{z}_t - \hat{\pi}_t\right) - \tilde{\tau}_t + \tilde{g}_t + \tilde{s}_t \tag{29}$$

return long term bond:

$$\hat{r}_{t,t+1}^m = \frac{\beta}{\Gamma}\rho_m \hat{p}_{t+1}^m - \hat{p}_t^m \tag{30}$$

no arbitrage condition:

$$\hat{r}_t = \hat{r}^m_{t,t+1} \tag{31}$$

monetary policy rule:

$$\hat{r}_t = \rho_r \hat{r}_{t-1} + (1-\rho_R)\left[\phi_\pi \hat{\pi} + \phi_Y\left(\hat{y}_t - \hat{y}^N_t\right)\right] + \sigma_R \epsilon_{R,t} \tag{32}$$

fiscal policy rule:

$$\hat{\tau}_t = \rho_\tau \hat{\tau}_{t-1} + (1-\rho_\tau)\,\phi_b b^m_{t-1} + \sigma_\tau \epsilon_{\tau,t} \tag{33}$$

government spending:

$$\hat{g}_t = \rho_g \hat{g}_{t-1} + \sigma_g \epsilon_{g,t} \tag{34}$$

technology shock:

$$\hat{z}_t = \rho_z \hat{z}_{t-1} + \sigma_z \epsilon_{z,t} \tag{35}$$

cost-push shock:

$$\hat{\mu}_t = \rho_\mu \hat{\mu}_{t-1} + \sigma_\mu \epsilon_{\mu,t} \tag{36}$$

term premia shock:

$$\hat{s}_t = \rho_s \hat{s}_{t-1} + \sigma_s \epsilon_{s,t} \tag{37}$$

Measurement equations:

$$\text{annual inflation} = 400\bar{\pi} + 100\left(\hat{\pi}_t + \hat{\pi}_{t-1} + \hat{\pi}_{t-2} + \hat{\pi}_{t-3}\right) \tag{38}$$

$$\frac{\text{primary deficits}}{\text{government debt}} = 400\left[\frac{\Gamma}{\bar{b}^m}\left(\hat{g}_t - \hat{\tau}_t\right) - \Gamma\bar{\pi}\frac{(1-1/\beta)}{\bar{b}^m}\hat{b}^m_{t-1}\right. \tag{39}$$

$$\left. +\Gamma\bar{\pi}\left(1-1/\beta\right)\left(\hat{y}_t - \hat{y}_{t-1} + \hat{z}_t + \hat{\pi}_t\right)\right]$$

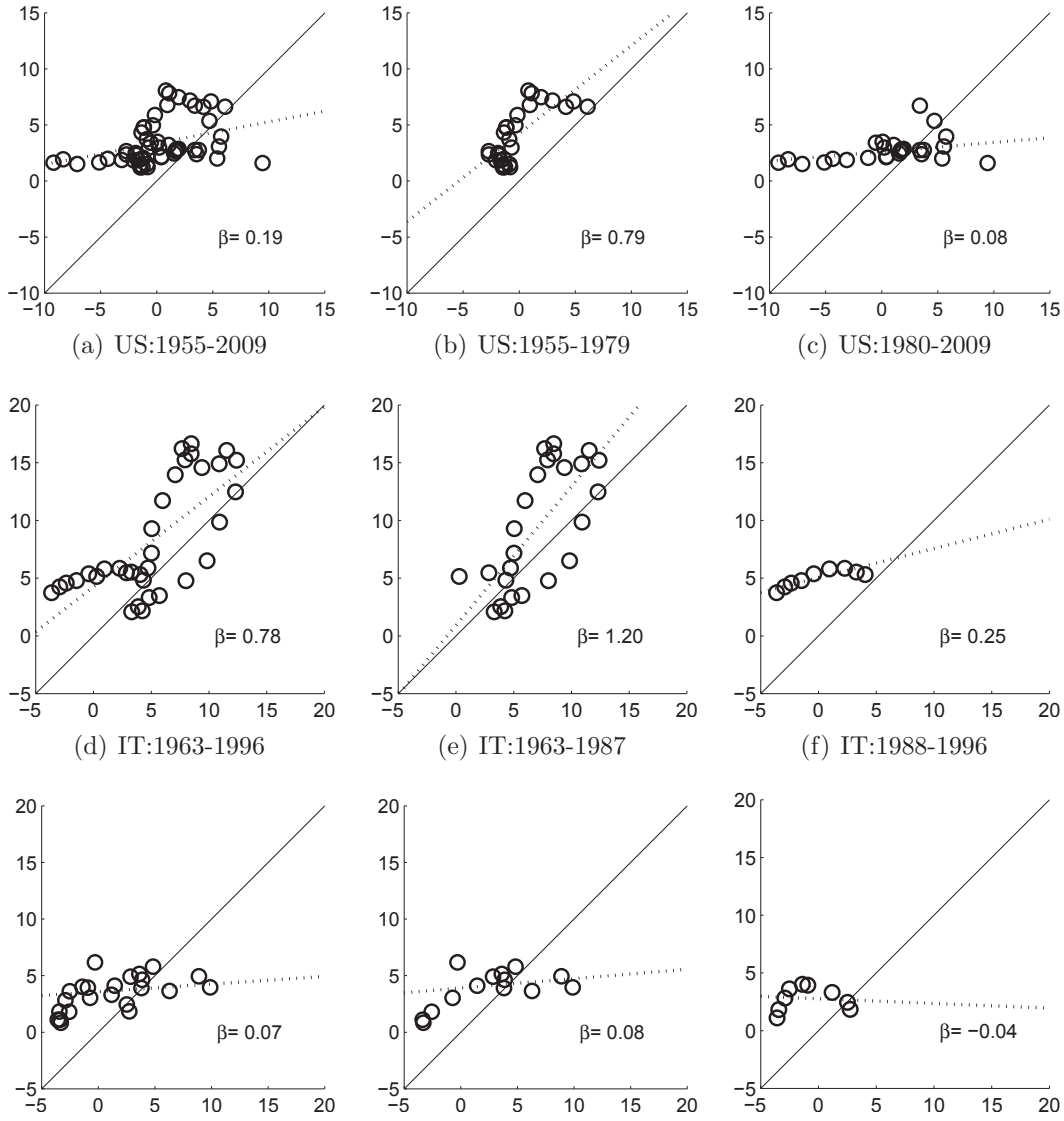
$$\text{real output growth} = 100\left(\gamma + \hat{y}_t - \hat{y}_{t-1} + \hat{z}_t\right) \tag{40}$$

$$\frac{\text{government purchases}}{\text{output}} = 100\left(\bar{g} + \hat{g}_t\right) \tag{41}$$

$$\frac{\text{government debt}}{\text{output}} = 100\left(4\bar{b}^m + \hat{b}^m_t\right) \tag{42}$$

$$\text{annualized interest rate} = 400\left[\left(\frac{\Gamma\bar{\pi}}{\beta} - 1\right) + \hat{r}_t\right] \tag{43}$$

# D Lucas Filter




(a) US:1955-2009 (b) US:1955-1979 (c) US:1980-2009

(d) IT:1963-1996 (e) IT:1963-1987 (f) IT:1988-1996

(g) GER:1963-1996 (h) GER:1963-1987 (i) GER:1988-1996

Figure 16: Scatter plots of filtered time series of inflation and deficits over debt. The dashed line indicates the slope of the scatter ($\beta$) and the solid line is the 45° line.