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Financial distress and fiscal inflation[☆]

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ABSTRACT

Is inflation 'always and everywhere a monetary phenomenon' or is it fundamentally a fiscal phenomenon? The answer hinges crucially on the underlying monetary–fiscal policy regime. Scant attention has been directed to the role of credit market frictions in discerning the policy regime, despite its growing importance in empirical macroeconomics. We augment a standard monetary model to incorporate fiscal details and credit market imperfections. These ingredients allow for both interpretations of the inflation process in a financially constrained environment. We find that introducing financial frictions to the model and adding financial variables to the dataset generate important identifying restrictions on the observed pattern between inflation and measures of financial and fiscal stress, to the extent that it overturns existing findings about which monetary–fiscal policy regime produced the U.S. data. To confront policy regime uncertainty, we propose the use of dynamic prediction pools and find strong cyclical patterns in the estimated historical regime weights.

1. Introduction

In any dynamic model with nominal government debt, there are two regions of the policy parameter space in which monetary and fiscal policies *jointly* determine inflation and stabilize debt. One region produces active monetary and passive fiscal policy or regime M, yielding the conventional monetarist/Wicksellian paradigm of inflation determination. Regime M assigns monetary policy to control inflation by raising the nominal interest rate aggressively with inflation and fiscal policy to stabilize debt by adjusting taxes or spending. The second region consists of passive monetary and active fiscal policy or regime F, producing the fiscal theory of the price level (Leeper, 1991; Woodford, 1995; Cochrane, 1999; Davig and Leeper, 2006; Sims, 2013). Under regime F, policy roles are reversed, with monetary policy responding weakly to inflation and fiscal instruments adjusting weakly to government debt.[1] Because these two policy regimes imply completely different mechanisms for price level determination and, therefore, starkly different policy advice, identifying the prevailing regime is a prerequisite to understanding the macroeconomy and to making good policy choices.

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[1] There is yet a third region of the policy parameter space that combines passive monetary and passive fiscal policy, exhibiting an indeterminate set of equilibria. See, for example, Clarida et al. (2000), Lubik and Schorfheide (2004), and Bhattarai et al. (2012, 2016).

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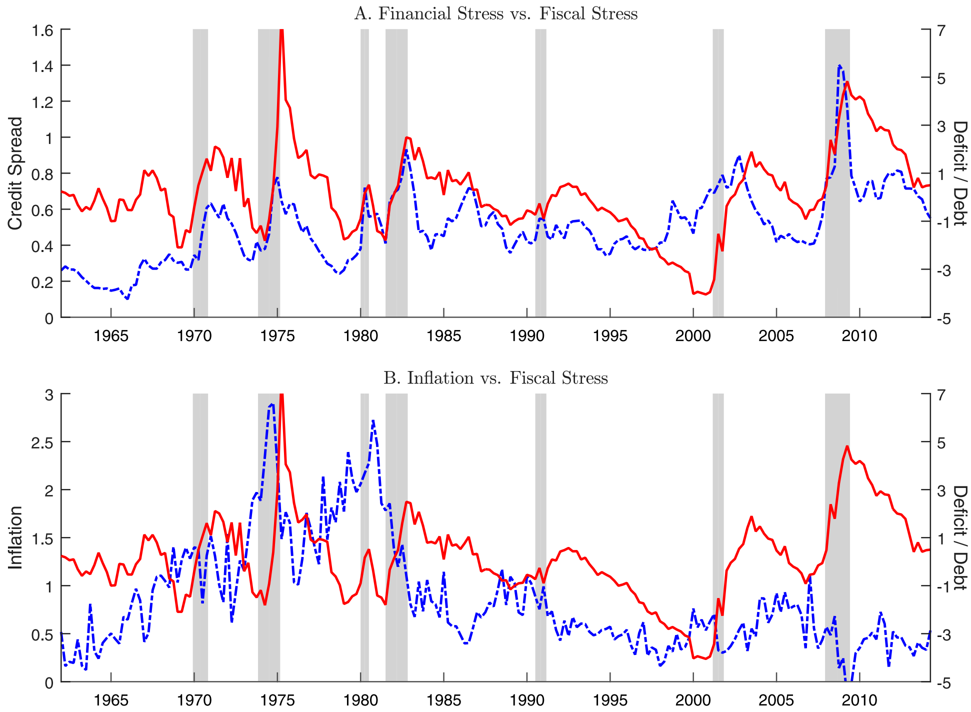

Fig. 1. Financial stress, fiscal stress, and inflation. Notes: The left and right vertical axes of Panel A measure the credit spread (blue dashed line, computed as Baa Corporate Bond Yield over Ten-Year Treasury Constant Maturity Rate) and the deficit-to-debt ratio (red solid line, constructed as in Sims (2011) by primary deficit as a proportion of lagged market value of privately held debt). The left and right vertical axes of Panel B measure the GDP deflator inflation (blue dashed line) and the deficit-to-debt ratio (red solid line). Shaded bars indicate recessions as designated by the National Bureau of Economic Research. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

While the popular surplus-debt regressions are subject to potential simultaneity bias that may produce misleading inferences about the nature of fiscal behavior, testing endeavors based on general equilibrium models with fixed policy rules, on the other hand, find nearly uniform statistical support for regime M over various subsamples of the precrisis U.S. data (Traum and Yang, 2011; Leeper and Li, 2017; Leeper et al., 2017). This consensus emerged even from periods of marked increases in fiscal stress (measured by deficit-to-debt ratio in Fig. 1), such as the mid-1970s, during which monetary policy appears to lose control over inflation.[2]

The literature on regime-switching monetary and fiscal policy constitutes a notable exception. By embedding the possibility of recurrent regime shifts in the formation of agents' expectations, this class of models uncovers alternating periods of regime M, the benign policy mix that many believe contributed to the steady decline in inflation in the 1980s, and regime F, the policy mix that some believe produced the run-ups in inflation in the 1970s, together with episodes of active (passive) monetary and active (passive) fiscal policy (Davig and Leeper, 2006; Bianchi, 2012, 2013; Bianchi and Ilut, 2017).[3]

Scant attention, however, has been paid to the empirical relevance of financial frictions in discerning the underlying policy regime despite its growing importance in macroeconomic modeling of business cycle fluctuations [Bernanke et al., 1999; Christiano et al., 2003, 2014; Del Negro et al., 2015, among others]. Such neglect is also surprising in light of the comovements between inflation and measures of financial and fiscal stress, as shown in Fig. 1—the Great Inflation from 1965 to 1982 concurred with several spikes in the credit spread and deficit-to-debt ratio, while the subsequent disinflation was associated with a downward trend in these variables, at least until the late 1990s. If inflation depends importantly on fiscal behavior and credit market conditions, a natural conjecture is that introducing financial frictions to the model and adding financial variables to the dataset may generate fresh identifying restrictions for the underlying policy regime.

This paper assesses the role of credit market imperfections in the identification of policy regimes. To that end, we extend a standard medium-scale dynamic stochastic general equilibrium (DSGE) model in two aspects. In particular, we follow Leeper et al. (2017) to fill in details of fiscal policy and incorporate the Bernanke et al. (1999) (henceforth, BGG) type of credit market frictions,

[2] Sims (2011) emphasizes that the primary surplus as a proportion of the total outstanding debt, which is closely related to our measure of fiscal stress, is the most natural single measure of fiscal stance.

[3] Chen et al. (2019) also allow monetary and fiscal policy to switch between active, passive, and optimal time-consistent rules. They find that optimizing policies fit the data to a degree comparable with the usual rule-based menu.

as in Christiano et al. (2014). These ingredients allow for a comprehensive study of monetary and fiscal policy interactions in a financially constrained environment.

Our key findings are threefold. First and foremost, adding financial frictions to the model and financial variables to the data improves the relative statistical fit of regime F. A rich set of model comparison exercises indicates that this result remains robust with respect to the sample period and fiscal details of the model and data. Surprisingly, regimes M and F become 'nearly' observationally equivalent in some cases, thereby overturning the clear-cut regime ranking found in the literature. As a byproduct of our model comparison exercises, we find that adding the deficit-to-debt ratio to the observables also improves the relative statistical fit of regime F, to the extent that it can fundamentally alter the regime ranking. This result largely corroborates that of Kliem et al. (2016a,b), who find that the low-frequency relationship between inflation and fiscal stance (defined as primary deficit over one-period lagged debt) plays a prominent role in discerning the underlying policy regime.

Second, the two policy regimes produce strikingly different inflation dynamics following a credit crunch. Contrary to regime M, which underlies the analysis of Christiano et al. (2014) and many others, elevated financial distress brings forth heightened fiscal uncertainty and inflation through the mechanism that regime F emphasizes. The implied comovement patterns between inflation, credit spread, and deficit-to-debt ratio under regime F turn out to be consistent with their major trends that Fig. 1 displays. More broadly, any exogenous disturbance that triggers different comovements in these variables across regimes, e.g., monetary and fiscal policy shocks, could potentially help to distinguish which policy regime generated the observed data.

Last, we consider an incomplete model space to confront policy regime uncertainty, where neither regime M nor regime F necessarily corresponds to the true model. By estimating a linear prediction pool model that dynamically combines both policy regimes, we find a strong cyclical pattern of transitions across regimes. In particular, despite the differences in model specification and dataset, the estimated historical regime weights exhibit pronounced cyclical fluctuations in all cases, with marked decreases in the relevance of regime M (or equivalently, sharp increases in the importance of regime F) accompanying the economic recessions.

Our paper complements the recent literature on the joint study of monetary and fiscal policy in the presence of financial frictions. For example, Cui (2016) develops a tractable macro model with endogenous asset liquidity to understand monetary–fiscal policy interactions with liquidity frictions. Xu and Serletis (2017) take a first pass at providing a theoretical demonstration of the fiscal theory of the price level in a world with borrowing constraints. Gomes and Seoane (2018) study the macroeconomic implications of each regime in a financially constrained environment and attribute the distinct postcrisis dynamics of the U.S. and Euro area to different policy arrangements. Li et al. (2018) show that a new Keynesian model with financial intermediaries and monetary–fiscal regime shifts can explain the time-varying correlation between returns on the market portfolio and Treasury bonds. To the best of our knowledge, this paper is the first to systematically assess the empirical relevance of financial and fiscal stress in the identification of policy regimes. While it is almost impossible to explore the entire model space, we perform extensive regime comparisons by estimating the marginal likelihoods for a total of 24 relevant models. Indeed, our analysis is rooted in the spirit of the Bayesian model scan framework proposed by Chib and Zeng (2019).

The rest of the paper is structured as follows. Section 2 outlines a standard medium-scale DSGE model augmented with financial frictions and a rich set of fiscal details for the empirical analysis. Section 3 reports the estimation results and explores the mechanism at work. Section 4 estimates a dynamic prediction pool model to confront policy regime uncertainty. Section 5 concludes.

2. Empirical models of policy interaction

This section outlines the benchmark new Keynesian model and its variants for the subsequent empirical analysis. We build two key features into an otherwise standard medium-scale DSGE model presented in Christiano et al. (2005) and Smets and Wouters (2007). In particular, we follow Leeper et al. (2017) to fill in details of fiscal policy and incorporate BGG-type credit market frictions as in Christiano et al. (2014). These ingredients allow for a comprehensive study of monetary and fiscal policy interactions in a financially constrained environment. A more detailed description of the model, steady state, and log-linearized equilibrium system can be found in the Online Appendix.

2.1. Firms

The production sector consists of firms that produce intermediate and final goods. A perfectly competitive final goods producer uses intermediate goods supplied by a continuum of intermediate goods producers indexed by i on the interval $[0,1]$ to produce the final goods. The production technology $Y_t \leq \left(\int_0^1 Y_t(i)^{1/(1+\eta_t^p)} di\right)^{1+\eta_t^p}$ is constant-return-to-scale. The price markup shock η_t^p follows the process $\ln\eta_t^p = (1-\rho_p)\ln\eta^p + \rho_p \ln\eta_{t-1}^p + \varepsilon_t^p$, where $\rho_p \in (0,1)$ and $\varepsilon_t^p \sim$ i.i.d.$\mathbb{N}(0,\sigma_p^2)$. η^p is the steady-state value of η_t^p. Y_t is the aggregate demand for the final goods. $Y_t(i)$ is the intermediate goods produced by firm i, which is the input demanded by the final goods producer.

Let $P_t(i)$ denote the price of intermediate goods and P_t the aggregate price index of the final goods. Therefore, we obtain the input demand function $Y_t(i) = \left(P_t(i)/P_t\right)^{-(1+\eta_t^p)/\eta_t^p} Y_t$, $\forall i$, and the relation $P_t = \left(\int_0^1 P_t(i)^{-1/\eta_t^p} di\right)^{-\eta_t^p}$.

Each intermediate goods producer follows a Cobb–Douglas production technology $Y_t(i) = K_t(i)^\alpha \left(A_t L_t^d(i)\right)^{1-\alpha} - A_t\Omega$, where $K_t(i)$ and $L_t^d(i)$ are the capital and the amount of 'packed' labor input rented by firm i at time t, and α is the income share of capital. A_t is the labor-augmenting neutral technology shock, and its growth rate $u_t^a \equiv \ln(A_t/A_{t-1})$ follows the process $u_t^a = (1-\rho_a)\gamma + \rho_a u_{t-1}^a + \varepsilon_t^a$, where $\rho_a \in (0,1)$ and $\varepsilon_t^a \sim$ i.i.d.$\mathbb{N}(0,\sigma_a^2)$. γ denotes the growth rate of A_t along the balanced growth path. The parameter Ω represents the fixed cost of production.

Intermediate goods producers maximize their profits in two stages. First, they take the input prices, i.e., nominal wage W_t and nominal rental rate of capital R_t^k, as given and rent $L_t^d(i)$ and $K_t(i)$ in perfectly competitive factor markets. Second, they choose the prices that maximize their discounted real profits. Here, we introduce a Calvo-pricing mechanism for nominal price rigidities. Specifically, a fraction $1-\omega_p$ of firms can change their prices each period. All other firms can only partially index their prices by the rule $P_t(i) = P_{t-1}(i)\left(\pi_{t-1}^{\chi_p}\pi^{1-\chi_p}\right)$, where $P_{t-1}(i)$ is indexed by the geometrically weighted average of past inflation π_{t-1} and steady-state inflation π. The weight $\chi_p \in [0,1]$ controls the degree of partial indexation. In a symmetric equilibrium, all the firms that can reoptimize their prices will choose the same price P_t^*. As a consequence, inflation evolves according to $1 = (1-\omega_p)(\pi_t^*)^{-1/\eta_t^p} + \omega_p\left[\left(\pi_{t-1}/\pi\right)^{\chi_p}\left(\pi/\pi_t\right)\right]^{-1/\eta_t^p}$, where $\pi_t^* \equiv P_t^*/P_t$.

2.2. Households

The economy is populated by a continuum of households indexed by j on the interval $[0,1]$. Each optimizing household j derives utility from real consumption $C_t(j)$ relative to a habit stock defined in terms of lagged aggregate consumption hC_{t-1}, where $h \in [0,1]$. Each household j supplies a continuum of differentiated labor services $L_t(j,l)$, where $l \in [0,1]$. The aggregate quantity of these labor services is $L_t(j) \equiv \int_0^1 L_t(j,l)dl$. Households maximize their expected lifetime utility $\mathbb{E}_0\sum_{t=0}^{\infty}\beta^t u_t^b\left[\ln(C_t(j)-hC_{t-1}) - L_t(j)^{1+\xi}/(1+\xi)\right]$, where β is the discount rate and ξ is the inverse of Frisch labor supply elasticity. u_t^b is an exogenous preference shock that follows the process $\ln u_t^b = (1-\rho_b)\ln u^b + \rho_b \ln u_{t-1}^b + \varepsilon_t^b$, where $\rho_b \in (0,1)$ and $\varepsilon_t^b \sim \text{i.i.d.}\mathbb{N}(0,\sigma_b^2)$. u^b is the steady-state value of u_t^b, which by assumption is 1.

Households make nominal deposits D_t at the financial intermediary, e.g., banks, that pay a noncontingent nominal gross interest rate R_t at time $t+1$. Households also have access to a portfolio of long-term nominal government bonds B_t, which sells at price P_t^B at time t. Following Woodford (2001), this general bond portfolio consists of perpetuities with coupons that decay exponentially. Specifically, we assume that a bond issued in period t pays ρ^k dollars $k+1$ periods later for each $k \geq 0$ with some decay factor $\rho \in [0,1]$. This assumption allows us to mimic the behavior of an arbitrary bond maturity structure with a single parameter ρ. A consol is the special case when $\rho = 1$, and a one-period bond arises when $\rho = 0$. In an environment with stable prices, the average maturity of such a bond portfolio is $(1-\rho\beta)^{-1}$.

The jth household's budget constraint is $P_tC_t(j) + P_t^B B_t(j) + D_t(j) = (1+\rho P_t^B)B_{t-1}(j) + R_{t-1}D_{t-1}(j) + \int_0^1 W_t(l)L_t(j,l)dl - P_tT_t(j) + P_t\Pi_t(j) + P_tTr_t(j)$, where $W_t(l)$ is the nominal wage charged by the household for type l labor service, $T_t(j)$ is the lump-sum tax net of transfer, $\Pi_t(j)$ is the profits from owning intermediate goods firms, and $Tr_t(j)$ is the net real transfer from entrepreneurs that will be discussed later.

We assume that the labor market is totally demand driven and that the demand for each type of labor l is uniformly allocated among households. Therefore, the total hours worked for each household is equal in equilibrium, i.e., $L_t(j) = L_t$. Under this assumption, we also have $L_t(j,l) = L_t(l)$ for all l.

There are perfectly competitive labor packers that hire a continuum of differentiated labor inputs $L_t(l)$, pack them to produce an aggregate labor service and then sell it to intermediate goods producers. The labor packer uses the Dixit–Stiglitz aggregator for labor aggregation $L_t^d = \left(\int_0^1 L_t(l)^{1/(1+\eta_t^w)}dl\right)^{1+\eta_t^w}$, where the wage markup shock η_t^w follows the process $\ln\eta_t^w = (1-\rho_w)\ln\eta^w + \rho_w\ln\eta_{t-1}^w + \varepsilon_t^w$, where $\rho_w \in (0,1)$ and $\varepsilon_t^w \sim \text{i.i.d.}\mathbb{N}(0,\sigma_w^2)$. η^w is the steady-state value of η_t^w. L_t^d is the aggregate labor service demanded by intermediate goods producers. $L_t(l)$ is the lth type of labor service supplied by all households and demanded by the labor packer.

Let $W_t(l)$ denote the differentiated nominal wage for type-l labor charged by the households and W_t the nominal wage index of the aggregate labor service. The labor packer's profit maximization yields the labor demand function $L_t(l) = L_t^d\left(W_t(l)/W_t\right)^{-(1+\eta_t^w)/\eta_t^w}$, $\forall l$, and the relation $W_t = \left(\int_0^1 W_t(l)^{-1/\eta_t^w}dl\right)^{-\eta_t^w}$.

For the optimal wage setting problem, we adopt the Calvo-pricing mechanism for nominal wage rigidities. Specifically, of all the types of labor services within each household, only the wages of a fraction $1-\omega_w$ can be changed each period. The wages for all other types of labor services follow a partial indexation rule $W_t(l) = W_{t-1}(l)\left(\pi_{t-1}e^{u_{t-1}^a}\right)^{\chi_w}(\pi e^{\gamma})^{1-\chi_w}$, where $W_{t-1}(l)$ is indexed by the geometrically weighted average of the growth rates of nominal wages in the past period and in the steady state, respectively. The growth rate of nominal wages is the product of the inflation rate and the growth rate of the real economy. The weight $\chi_w \in [0,1]$ controls the degree of partial indexation. In a symmetric equilibrium, all the reoptimized wages are equal to W_t^*. As a consequence, the aggregate real wage index w_t evolves according to $w_t^{-1/\eta_t^w} = (1-\omega_w)(w_t^*)^{-1/\eta_t^w} + \omega_w\left[\left(\pi_{t-1}e^{u_{t-1}^a}\right)^{\chi_w}/(\pi e^{\gamma})^{\chi_w}\cdot(\pi e^{\gamma})/(\pi_t e^{u_t^a})\cdot w_{t-1}\right]^{-1/\eta_t^w}$, where the real wage is detrended by A_t for stationarity, i.e., $w_t \equiv W_t/(P_tA_t)$ and $w_t^* \equiv W_t^*/(P_tA_t)$.

2.3. Capital goods producers

There are perfectly competitive capital goods producers who buy installed capital X_t from entrepreneurs for the nominal price Q_t^k and add new investment I_t using the final goods to generate new installed capital X_{t+1} for the next period, which they sell back to entrepreneurs for the nominal price Q_t^k at the end of period t. Note that the law of motion is given by $X_{t+1} = X_t + \tilde{u}_t^i\left(1-S\left(I_t/I_{t-1}\right)\right)I_t$, where $S(\cdot)I_t$ is an investment adjustment cost. Specifically, we assume a functional form $S(I_t/I_{t-1}) = \frac{s}{2}[I_t/I_{t-1} - e^{\gamma}]^2$ with the property that $S(e^{\gamma}) = S'(e^{\gamma}) = 0$ and $S''(e^{\gamma}) = s > 0$. Therefore, a real cost is incurred when the short-run growth rate of investment deviates from its long-run growth rate. In the capital accumulation process, there is an investment-specific efficiency shock $\tilde{u}^i_t$ that follows the process $\ln\tilde{u}^i_t = (1-\rho_i)\ln\tilde{u}^i + \rho_i \ln\tilde{u}^i_{t-1} + \tilde{\varepsilon}^i_t$, where $\rho_i \in (0,1)$ and $\tilde{\varepsilon}^i_t \sim$ i.i.d.$\mathbb{N}(0, \tilde{\sigma}^2_i)$. $\tilde{u}^i$ is the steady-state value of $\tilde{u}^i_t$.

By imposing market clearing, i.e., $X_t = (1-\delta)\bar{K}_{t-1}$ and $X_{t+1} = \bar{K}_t$, the law of motion for capital can be rewritten as $\bar{K}_t = (1-\delta)\bar{K}_{t-1} + \tilde{u}^i_t \left(1 - S\left(I_t/I_{t-1}\right)\right) I_t$, with the capital depreciation rate δ.

2.4. Entrepreneurs

There is a continuum of entrepreneurs indexed by e. Each entrepreneur purchases installed capital $\bar{K}_{t-1}(e)$ from the capital goods producers for price Q^k_{t-1} at the end of period $t-1$ using her own net worth $N_{t-1}(e)$ and a loan $B^d_{t-1}(e)$ from the financial intermediary, e.g., banks. The balance sheet of entrepreneurs in nominal terms implies that $Q^k_{t-1}\bar{K}_{t-1}(e) = B^d_{t-1}(e) + N_{t-1}(e)$. In period t, the entrepreneur rents capital out to intermediate goods producers, earning a nominal rental rate R^k_t per unit of effective capital, which is the installed capital multiplied by a utilization rate. Additionally, in period t, each entrepreneur is subject to an idiosyncratic shock $\omega_t(e)$ that converts the installed capital $\bar{K}_{t-1}(e)$ into efficiency units, i.e., $\omega_t(e)\bar{K}_{t-1}(e)$. The shock is i.i.d. across entrepreneurs and over time and may increase or shrink each entrepreneur's capital. As in BGG, we assume that the shock $\omega_t(e)$ follows a log-normal distribution with unit mean. Specifically, $\ln\omega_t(e) \sim N(\mu_{\omega,t-1}, \sigma^2_{\omega,t-1})$ where $\mu_{\omega,t-1}$ is such that $\mathbb{E}\omega_t(e) = 1$. Since a log-normal random variable has $\mathbb{E}\omega_t(e) = \exp\left\{\mu_{\omega,t-1} + \frac{1}{2}\sigma^2_{\omega,t-1}\right\}$, it must be the case that $\mu_{\omega,t-1} = -\frac{1}{2}\sigma^2_{\omega,t-1}$. As in Christiano et al. (2014), we call $\sigma_{\omega,t}$ a risk shock that characterizes the extent of cross-sectional dispersion in ω. Here, we allow $\sigma_{\omega,t}$ to be time varying according to the process $\ln\sigma_{\omega,t} = (1-\rho_{\sigma_\omega})\ln\sigma_\omega + \rho_{\sigma_\omega}\ln\sigma_{\omega,t-1} + \tilde{\varepsilon}^{\sigma_\omega}_t$, where $\rho_{\sigma_\omega} \in (0,1)$ and $\tilde{\varepsilon}^{\sigma_\omega}_t \sim$ i.i.d.$\mathbb{N}(0, \tilde{\sigma}^2_{\sigma_\omega})$. σ_ω is the steady-state value of $\sigma_{\omega,t}$. Let $F_{t-1}(\omega)$ denote the CDF function of ω at time t. It is clear from the timing that the distribution needs to be known at time $t-1$ and the shock is realized at time t. After observing the shock at time t, the entrepreneur may choose a level of capital utilization $u_t(e)$, which incurs a real cost in terms of consumption goods, i.e., $a(u_t(e))$ per unit of capital. At the end of period t, the entrepreneur sells her undepreciated capital back to the capital goods producers.

In period t, the entrepreneur's revenue contains rental revenue and gains from the capital sale net of capital utilization cost, which in nominal terms is $\left\{R^k_t u_t(e) + (1-\delta)Q^k_t - P_t a(u_t(e))\right\}\omega_t(e)\bar{K}_{t-1}(e)$. It can be written equivalently as $\omega_t(e)\tilde{R}^k_t(e)Q^k_{t-1}\bar{K}_{t-1}(e)$, where we define the gross nominal return to capital for entrepreneurs as $\tilde{R}^k_t(e) \equiv [R^k_t u_t(e) + (1-\delta)Q^k_t - P_t a(u_t(e))]/Q^k_{t-1}$. It can be shown that the entrepreneur's choice of utilization rate does not depend on index e. Hence, we may also drop the index for $\tilde{R}^k_t$, which, in log-linearized form, is

$$\hat{\tilde{R}}^k_t = \frac{r^k}{r^k + 1 - \delta}\hat{r}^k_t + \frac{1-\delta}{r^k + 1 - \delta}\hat{q}_t - \hat{q}_{t-1} + \hat{\pi}_t \tag{2.1}$$

where $r^k_t \equiv R^k_t/P_t$, $q_t \equiv Q^k_t/P_t$, and r^k is the steady-state value of r^k_t. For $a(u)$, we specifically assume $a(u) = \gamma_1(u-1) + \frac{\gamma_2}{2}(u-1)^2$, where $\gamma_1, \gamma_2 \geq 0$. In steady state, $u = 1$ so $a(1) = 0$. It can be shown that $a'(1) = \gamma_1$ and $a''(1) = \gamma_2$. As in Leeper et al. (2017), we define a parameter $\psi \in [0,1)$ such that $\frac{a''(1)}{a'(1)} \equiv \frac{\psi}{1-\psi}$, implying $\psi = \frac{\gamma_2}{\gamma_1+\gamma_2}$. As $\psi \to 1$, utilization costs become infinite, and the capital utilization rate becomes constant.

In period $t-1$, the debt contract between the entrepreneur and the financial intermediary consists of $(B^d_{t-1}(e), R^d_t(e))$, where $R^d_t(e)$ is the interest rate for the contract. Let $\bar{\omega}_t(e)$ denote the threshold level of $\omega_t(e)$ below which the entrepreneur cannot pay back loans so will default and declare bankruptcy. Specifically, $\bar{\omega}_t(e)$ is defined by $\bar{\omega}_t(e)\tilde{R}^k_t Q^k_{t-1}\bar{K}_{t-1}(e) = R^d_t(e)B^d_{t-1}(e)$. In period t, for $\omega_t(e) < \bar{\omega}_t(e)$, the financial intermediary will monitor the entrepreneur and extract a fraction $(1-\mu^e)$ of her revenue $\tilde{R}^k_t Q^k_{t-1}\bar{K}_{t-1}(e)$, and μ^e represents the fraction of bankruptcy costs. This mechanism, as proposed by BGG, captures the asymmetries of information between lenders and borrowers and the need to have costly state verification. In such an environment, banks can observe the realization of $\omega_t(e)$ only by undertaking costly monitoring and will therefore include a premium in the interest rate on loans to protect themselves against default risk. It can be shown in log-linearized form that the expected capital return premium, i.e., the credit spread, fluctuates with entrepreneurs' leverage and riskiness according to

$$\mathbb{E}_t[\hat{\tilde{R}}^k_{t+1} - \hat{R}_t] = \zeta_{sp,b}(\hat{q}_t + \hat{\bar{k}}_t - \hat{n}_t) + \zeta_{sp,\sigma_\omega}\hat{\sigma}_{\omega,t} \tag{2.2}$$

where $\bar{k}_t \equiv \bar{K}_t/A_t$, $n_t \equiv N_t/(P_t A_t)$, and $(\zeta_{sp,b}, \zeta_{sp,\sigma_\omega})$ are the elasticities with respect to leverage and risk. Note that if $\zeta_{sp,b} = 0$ and $\hat{\sigma}_{\omega,t} = 0$, financial frictions dissipate and (2.1) and (2.2) reduce to the familiar arbitrage condition between the return to capital and the risk-free rate.

After the entrepreneurs have sold their undepreciated capital, collected capital rental receipts and settled their obligations to their loans at the end of period t, a fraction of each entrepreneur's assets is transferred to households. The rest remains in the hand of the entrepreneurs. In addition, each entrepreneur receives a lump-sum transfer from the households.

2.5. Monetary and fiscal policy

The central bank implements monetary policy according to an interest rate rule. It reacts to deviations of inflation and output from their steady states

$$\hat{R}_t = \rho_r\hat{R}_{t-1} + (1-\rho_r)(\phi_\pi\hat{\pi}_t + \phi_y\hat{y}_t) + u^m_t \tag{2.3}$$

where ρ_r determines the degree of policy smoothing, (ϕ_π, ϕ_y) controls the responsiveness to inflation and output deviations, $y_t \equiv Y_t/A_t$, and u^m_t is a monetary policy shock that follows the process $u^m_t = \rho_{em}u^m_{t-1} + \varepsilon^m_t$, where $\rho_{em} \in (0,1)$ and $\varepsilon^m_t \sim$ i.i.d.$\mathbb{N}(0, \sigma^2_m)$.

The government collects a lump-sum tax net of transfer, T_t, and sells the nominal bond portfolio, B_t, to finance its interest payments and consumption expenditures, G_t. Therefore, fiscal choices satisfy the government budget constraint $P_t^B B_t + P_t T_t = (1 + \rho P_t^B)B_{t-1} + P_t G_t + TP_t$, where TP_t is a shock that captures a set of features that are not explicitly modeled, such as changes in the maturity structure and the term premium. Rewriting the variables in terms of the ratio to GDP, we express the government budget constraint as

$$b_t^s + \tau_t^s = \frac{R_{t-1,t}^B b_{t-1}^s}{\pi_t \Delta_t^y} + g_t^s + tp_t \tag{2.4}$$

where $b_t^s \equiv P_t^B B_t/(P_t Y_t)$, $\tau_t^s \equiv T_t/Y_t$, $g_t^s \equiv G_t/Y_t$, and $tp_t \equiv TP_t/(P_t Y_t)$. $\Delta_t^y \equiv Y_t/Y_{t-1}$ is the gross growth rate of output. Note that $R_{t-1,t}^B \equiv (1 + \rho P_t^B)/P_{t-1}^B$ is the realized return of the long-term bond portfolio. The shock tp_t follows the process $tp_t = \rho_{tp} tp_{t-1} + \epsilon_t^{tp}$, where $\rho_{tp} \in (0, 1)$ and $\epsilon_t^{tp} \sim$ i.i.d.$\mathbb{N}(0, \sigma_{tp}^2)$. In what follows, we denote $(\hat{b}_t^s, \hat{\tau}_t^s, \hat{g}_t^s)$ the level of deviations of these variables from their steady states, i.e., $\hat{x}_t \equiv x_t - x$ for $x \in \{b^s, \tau^s, g^s\}$, to avoid having the percentage change of a percentage.

Fiscal variables, as indicated in Fig. 1, are persistent and variable. To capture these features, we follow Sims (2012) and embed a 'long-run risk' type of component into a possibly countercyclical tax rule

$$\hat{\tau}_t^s = \gamma_b^\tau \hat{b}_{t-1}^s + \gamma_y^\tau \hat{y}_t + u_{\tau,t}^l + u_{\tau,t}^s \tag{2.5}$$

where $(\gamma_b^\tau, \gamma_y^\tau)$ controls the responsiveness to deviations of lagged debt and current output. The tax shock contains both a very persistent long-run component $u_{\tau,t}^l$ with small variance to capture the low frequency movement in fiscal variables and a mildly persistent short-run component $u_{\tau,t}^s$. Their processes are $u_{\tau,t}^l = \rho_{el} u_{\tau,t-1}^l + \epsilon_{\tau,t}^l$ and $u_{\tau,t}^s = \rho_{e\tau} u_{\tau,t-1}^s + \epsilon_{\tau,t}^s$, where $\rho_{el}, \rho_{e\tau} \in (0, 1)$. Moreover, government consumption follows the rule

$$\hat{g}_t^s = \rho_g \hat{g}_{t-1}^s - (1 - \rho_g)(\gamma_b^g \hat{b}_{t-1}^s + \gamma_y^g \hat{y}_t) + u_{g,t} \tag{2.6}$$

where ρ_g measures the degree of policy smoothing, (γ_b^g, γ_y^g) controls the responsiveness to deviations of lagged debt and current output, and $u_{g,t}$ is a fiscal spending shock that follows the process $u_{g,t} = \rho_{eg} u_{g,t-1} + \epsilon_{g,t}$, where $\rho_{eg} \in (0, 1)$. The presence of response to output deviations in (2.5) and (2.6) renders fiscal instruments automatically stabilizing, which plays a similar role as Leeper et al. (2017)'s distorting steady-state taxes. Note that all the innovations to fiscal shocks $\{\epsilon_{\tau,t}^l, \epsilon_{\tau,t}^s, \epsilon_{g,t}\} \sim$ i.i.d. normal with mean 0 and standard deviation $\{\sigma_\tau^l, \sigma_\tau^s, \sigma_g\}$, respectively.

There are two regions of the policy parameter space corresponding to regimes M and F.[4] The fundamental difference between the two policy regimes lies in their distinct fiscal financing schemes. We highlight all financing possibilities of government debt that can be gleaned from the linearized version of government budget constraints (2.4)

$$\hat{b}_t^s = -\underbrace{(\hat{\tau}_t^s - \hat{g}_t^s)}_{\text{primary surplus}} - \underbrace{\frac{b^s}{\beta}\hat{\pi}_t}_{\text{surprise inflation}} + \underbrace{\frac{b^s \rho}{\pi e^\gamma}\hat{P}_{B,t}}_{\text{bond price}} - \underbrace{\frac{b^s}{\beta}\hat{\Delta}_t^y}_{\text{output growth}} + \underbrace{\frac{1}{\beta}(\hat{b}_{t-1}^s - b^s \hat{P}_{B,t-1})}_{\text{predetermined term}} + \underbrace{tp_t}_{\text{shock}}$$

which makes it clear that fiscal consolidation can be accomplished through several channels – higher primary surplus, surprise inflation, lower bond price, and higher output growth – or any of their combinations, regardless of the policy regime in place. In particular, while regime M relies primarily on direct taxation, regime F hinges crucially on the debt revaluation effects of higher inflation and lower bond prices.

2.6. Model variants

In addition to the benchmark specification ($\mathcal{M}_1$) above, we also consider two variants of $\mathcal{M}_1$ to check the robustness of our results, including the original model of Leeper et al. (2017) with distorting taxes and government consumption valued as a public good ($\mathcal{M}_2$) as well as its lump-sum taxes version with automatic stabilizers ($\mathcal{M}_3$). Specifically,

- In $\mathcal{M}_2$, the government collects revenues from capital, labor, and consumption taxes and sells a nominal bond portfolio to finance its interest payments and expenditures. The fiscal choices satisfy the flow budget constraint $P_t^B B_t + \tau_t^K R_t^k K_t + \tau_t^L W_t L_t + P_t \tau_t^C C_t = (1 + \rho P_t^B)B_{t-1} + P_t G_t + P_t Z_t + TP_t$, where τ^K, τ^L, and τ^C are the distorting tax rates on capital income, labor income and consumption sales, and Z_t is lump-sum transfer. As in Leeper et al. (2017), the fiscal instruments follow simple rules $\hat{g}_t = \rho_g \hat{g}_{t-1} - (1 - \rho_g)\gamma_g \hat{b}_{t-1}^s + u_t^g$ and $\hat{z}_t = \rho_z \hat{z}_{t-1} - (1 - \rho_z)\gamma_z \hat{b}_{t-1}^s + u_t^z$, where $g_t = G_t/A_t$, $z_t = Z_t/A_t$, and $\rho_g, \rho_z \in [0, 1)$ are typically close to unity since fiscal variables are generally quite persistent in the data. The fiscal policy shock u_t^f, $f \in \{g, z\}$, follows AR process $u_t^f = \rho_{ef} u_{t-1}^f + \epsilon_t^f$, where $\rho_{ef} \in [0, 1)$ and $\epsilon_t^f \sim \mathbb{N}(0, \sigma_f^2)$.
- In $\mathcal{M}_3$, we replace distorting taxes in $\mathcal{M}_2$ with lump-sum taxes and allow fiscal instruments to contain an automatic stabilizer that responds to the output gap.[5] The fiscal choices satisfy the flow budget constraint $P_t^B B_t + P_t T_t = (1 + \rho P_t^B)B_{t-1} + P_t G_t + TP_t$, where T_t is a lump-sum tax net of transfer. The fiscal instruments follow simple rules $\hat{\tau}_t = \rho_\tau \hat{\tau}_{t-1} + (1 - \rho_\tau)(\gamma_b^\tau \hat{b}_{t-1}^s + \gamma_y^\tau \hat{y}_t) + u_t^\tau$ and $\hat{g}_t = \rho_g \hat{g}_{t-1} - (1 - \rho_g)(\gamma_b^g \hat{b}_{t-1}^s + \gamma_y^g \hat{y}_t) + u_t^g$, where $\tau_t = T_t/A_t$ and $\rho_\tau, \rho_g \in [0, 1)$ are again close to unity given the high persistence in the fiscal data. The fiscal policy shock u_t^f, $f = \{\tau, g\}$, follows AR process $u_t^f = \rho_{ef} u_{t-1}^f + \epsilon_t^f$, where $\rho_{ef} \in [0, 1)$ and $\epsilon_t^f \sim \mathbb{N}(0, \sigma_f^2)$.

[4] In the Online Appendix, we present an analytical characterization of the key features under each regime using a simple frictionless model of inflation determination.

[5] The automatic stabilizer specified here plays a similar role as the steady-state distortionary taxes in Leeper et al. (2017).

Table 1
Calibrated parameters.

Average duration of government debt	AD	20
Discount factor	β	0.99
Depreciation rate	δ	0.025
Capital income share	α	0.33
Price markup	η^p	0.14
Wage markup	η^w	0.14
Government expenditure to GDP ratio	g/y	0.11
Government debt to GDP ratio	b/y	1.47
Steady-state probability of entrepreneurial default	$\bar{F}$	0.03
Fraction of entrepreneur's remained assets	γ^e	0.99
Lag coefficient of long-run tax component	ρ_{el}	0.99
Standard deviation of long-run tax shock	$100\sigma_\tau^l$	0.018

To evaluate the role of credit market imperfections in the identification of policy regimes, we also consider the three models ($\mathcal{M}_1$–$\mathcal{M}_3$) when financial frictions dissipate, i.e., $\zeta_{sp,b} = 0$ and $\hat{\sigma}_{\omega,t} = 0$. In this case, (2.1) and (2.2) collapse to the familiar arbitrage condition between the return to capital and the risk-free rate. All models are log-linearized around the steady state and solved using Sims' (2002) procedure.

3. Regime comparison

We apply Bayesian methods to estimate each regime-dependent model over two subsamples: the pre-Volcker era, 1962: Q1–1979: Q2; and the post-Volcker era, 1984: Q1–2007: Q4.[6] The common set of quarterly observables includes log differences of private consumption, investment, real wage, and GDP deflator; log hours worked; and the federal funds rate. In addition, we add two fiscal variables, i.e., government consumption and real market value of debt, to the dataset: ratios to GDP for $\mathcal{M}_1$ and log differences for $\mathcal{M}_2$ and $\mathcal{M}_3$.

As pointed out by Kliem et al. (2016a,b), the low-frequency relationship between inflation and fiscal stance (defined as primary deficit over one-period lagged debt) also plays an important role in discerning the underlying policy regime. To assess the empirical relevance of financial and fiscal stress in regime identification, we allow either the credit spread, deficit-to-debt ratio, or both to enter or not enter the list of observables for $\mathcal{M}_1$–$\mathcal{M}_3$. The credit spread is included in the data whenever the model incorporates financial frictions. Otherwise, this measure of financial stress is excluded from the data. The resulting model space contains a total of 24 relevant models that are simultaneously confronted with the data. See the Online Appendix for details on data construction and measurement equations.

3.1. Prior

We first fix a few parameters for all model specifications in the estimation, which are shown in Table 1. The average duration of government bonds is assumed to be five years, i.e., $AD = 20$. The discount factor is $\beta = 0.99$. The quarterly depreciation rate for private capital, δ, is 0.025 so that the annual rate is 10 percent. The capital income share of total output is $\alpha = 0.33$. The price and wage markups are $\eta^p = \eta^w = 0.14$. Steady-state fiscal variables are calibrated to the mean values of the U.S. data. Federal government expenditure as a share of model-consistent output, i.e., GDP excluding net exports, is 0.11, and the ratio of federal debt to model-consistent output is 1.47. For those model specifications with financial frictions, we follow Del Negro et al. (2015) and fix the steady-state probability of entrepreneurial default $\bar{F}$ at 0.03 and the fraction of entrepreneurs' remaining assets γ^e at 0.99.

Let θ be a vector collecting all estimated parameters and $p(\theta)$ its joint prior distribution. Table 2 lists the marginal prior distributions, which are assumed to be independent so that their product equals $p(\theta)$.[7] Our prior settings are broadly consistent with the literature, e.g., Del Negro et al. (2015), Bianchi and Ilut (2017), and Leeper et al. (2017). The only exceptions are the priors for the nominal rigidity parameters, i.e., $(\omega_p, \omega_w, \chi_p, \chi_w)$, which are somewhat tighter than those assumed in the literature. Specifically, they follow beta distributions with a standard deviation of 0.05. The relatively informed priors serve as regularization devices for facilitating posterior mode-finding and sampling.[8]

To reflect the two distinct policy regimes, we specify two sets of priors on the policy parameters, each of which places almost all probability mass on the region of the parameter space that delivers a unique model solution consistent with a particular regime.[9] For monetary policy response to inflation, ϕ_π has a normal (beta) prior with mean 2 (0.5) and standard deviation 0.2 (0.1) under regime M (F). Regarding the fiscal policy parameters,

[6] Our full sample begins when the spread data first becomes available and ends before the federal funds rate nearly hit its effective lower bound.

[7] These priors are common for all models except for the fiscal policy parameters that will be discussed later.

[8] We have also experimented with a standard deviation of 0.2, as in Leeper et al. (2017), for the priors of these parameters but found that the posterior sampler often became stuck near the boundary of the determinacy regions.

[9] The boundary separating different regimes is not analytically available for sophisticated DSGE models.

Table 2
Prior distributions.

Parameter	Distribution	Para(1)	Para(2)	Parameter	Distribution	Para(1)	Para(2)
100γ	$\mathbb{N}$	0.4	0.05	ρ_a	$\mathbb{B}$	0.5	0.2
ξ	$\mathbb{G}$	2	0.5	ρ_b	$\mathbb{B}$	0.5	0.2
h	$\mathbb{B}$	0.5	0.2	ρ_i	$\mathbb{B}$	0.5	0.2
ψ	$\mathbb{B}$	0.6	0.15	ρ_p	$\mathbb{B}$	0.5	0.2
s	$\mathbb{G}$	6	1.5	ρ_w	$\mathbb{B}$	0.5	0.2
ω_p	$\mathbb{B}$	0.6	0.05	ρ_{tp}	$\mathbb{B}$	0.5	0.2
ω_w	$\mathbb{B}$	0.6	0.05	ρ_{em}	$\mathbb{B}$	0.5	0.15
χ_p	$\mathbb{B}$	0.5	0.05	$\rho_{e\tau}$	$\mathbb{B}$	0.5	0.15
χ_w	$\mathbb{B}$	0.5	0.05	ρ_{eg}	$\mathbb{B}$	0.5	0.15
ϕ_π, regime-M	$\mathbb{N}$	2	0.2	$100\sigma_a$	$\mathbb{IG}$-1	2	0.1
ϕ_π, regime-F	$\mathbb{B}$	0.5	0.1	$100\sigma_b$	$\mathbb{IG}$-1	2	0.1
ϕ_y	$\mathbb{N}$	0.125	0.05	$100\sigma_i$	$\mathbb{IG}$-1	2	0.1
ρ_r	$\mathbb{B}$	0.5	0.2	$100\sigma_p$	$\mathbb{IG}$-1	2	0.1
γ_b^τ, regime-M	$\mathbb{N}$	0.05	0.01	$100\sigma_w$	$\mathbb{IG}$-1	2	0.1
γ_b^g, regime-M	$\mathbb{N}$	0.05	0.01	$100\sigma_{tp}$	$\mathbb{IG}$-1	2	0.1
γ_b^τ, regime-F	–	0	0	$100\sigma_m$	$\mathbb{IG}$-1	2	0.1
γ_b^g, regime-F	–	0	0	$100\sigma_\tau^s$	$\mathbb{IG}$-1	2	0.1
γ_y^τ	$\mathbb{N}$	0.4	0.2	$100\sigma_g$	$\mathbb{IG}$-1	2	0.1
γ_y^g	$\mathbb{N}$	0.4	0.2	SP	$\mathbb{G}$	0.5	0.1
ρ_g	$\mathbb{B}$	0.5	0.2	$\zeta_{sp,b}$	$\mathbb{B}$	0.05	0.005
$\bar{L}$	$\mathbb{N}$	468	5	ρ_{σ_ω}	$\mathbb{B}$	0.75	0.15
$\bar{\pi}$	$\mathbb{G}$	0.75	0.25	$100\sigma_{\sigma_\omega}$	$\mathbb{IG}$-1	4	0.05

[1] Para (1) and Para (2) correspond to the means and standard deviations for Gamma ($\mathbb{G}$), Normal ($\mathbb{N}$), and Beta ($\mathbb{B}$) distributions; ν and s for the Inverse-Gamma Type-I ($\mathbb{IG}$-1) distribution, where $p(\sigma) \propto \sigma^{-\nu-1}\exp(-\frac{\nu s^2}{2\sigma^2})$.
[2] The effective prior is truncated at the boundary of the determinacy region.

- In $\mathcal{M}_1$, we fix the responses of fiscal instruments to government debt, i.e., γ_b^τ and γ_b^g, at zero under regime F but estimate these parameters under regime M. For both policy regimes, we set the lag coefficient of the long-run tax component ρ_{el} to be 0.99. Following Sims (2012), we assume the long-run tax component has a very small variance and set the standard deviation of long-run tax shock σ_τ^l to 0.018, which is one-tenth of the prior mean for the standard deviation of other structural shocks.
- In $\mathcal{M}_2$, we calibrate the steady-state distorting tax rates as in Leeper et al. (2017): the average federal labor tax rate τ^L is 0.186, the capital tax rate τ^K is 0.218, and the consumption tax rate τ^C is 0.023. Under regime M, we set $\rho_z = 0.98$ and $\rho_{ez} = 0.8$ and estimate the responses of fiscal instruments to government debt, i.e., γ_g and γ_z. Under regime F, we estimate ρ_z and ρ_{ez} and fix γ_g and γ_z at zero.
- In $\mathcal{M}_3$, we set $\rho_\tau = 0.98$ and $\rho_{e\tau} = 0.8$ under regime M and estimate them under regime F. As in $\mathcal{M}_1$, we fix γ_b^τ and γ_b^g at zero under regime F and estimate them under regime M.

3.2. Posterior estimates

The prior distribution $p(\theta)$ summarizes the researcher's initial views of the model parameters. This prior information is updated with the data Y via Bayes' theorem

$$p(\theta|Y) \propto p(Y|\theta)p(\theta) \tag{3.1}$$

where $p(Y|\theta)$ is the likelihood function and the posterior distribution $p(\theta|Y)$ characterizing the researcher's updated parameter beliefs is calculated up to the normalization constant (i.e., the marginal likelihood $p(Y)$). It is well known that the surface of $p(\theta|Y)$ can be highly irregular for high-dimensional DSGE models. To overcome the irregularity problems, we simulate the model parameters from their joint posterior distribution using the tailored randomized block Metropolis–Hastings (TaRB-MH) algorithm that was originally proposed by Chib and Ramamurthy (2010) and recently optimized as a MATLAB toolbox by Chib et al. (2021). For each regime-dependent model, we sample a total of 11,000 parameter draws and discard the first 1,000 draws as the burn-in phase. The resulting 10,000 draws form the basis for performing our posterior inference. Due to the efficiency gains achieved by the TaRB-MH algorithm, the number of draws is substantially smaller than that typically used for the conventional random-walk MH algorithm. The overall inefficiency factor is on average very low, with most values below 30. In conjunction with a rejection rate of approximately 50% in the MH step for each model, the small inefficiency factors suggest that the Markov chain mixes well.

Table 3 reports the posterior parameter means and 90% credible intervals for $\mathcal{M}_1$ with financial frictions.[10] The 90% credible intervals show that the posterior distributions for most parameters are different from the priors. Thus, the data are overall informative

[10] To conserve space, the posterior parameter estimates for all remaining cases are relegated to the Online Appendix.

Table 3
Posterior estimates, $\mathcal{M}_1$ with financial and fiscal stress.

	1962:Q1–1979:Q2				1984:Q1–2007:Q4			
	Regime-M		Regime-F		Regime-M		Regime-F	
	Mean	[5,95]	Mean	[5,95]	Mean	[5,95]	Mean	[5,95]
100γ	0.28	[0.20, 0.36]	0.31	[0.24, 0.39]	0.36	[0.29, 0.43]	0.36	[0.30, 0.42]
ξ	1.85	[1.15, 2.48]	1.89	[1.16, 2.57]	2.60	[1.72, 3.36]	2.08	[1.34, 2.79]
h	0.92	[0.89, 0.95]	0.95	[0.94, 0.97]	0.96	[0.94, 0.99]	0.95	[0.92, 0.97]
ψ	0.68	[0.45, 0.90]	0.57	[0.41, 0.72]	0.51	[0.27, 0.71]	0.53	[0.35, 0.70]
s	3.20	[2.24, 4.24]	2.72	[1.82, 3.56]	4.60	[3.15, 6.11]	4.39	[3.08, 5.57]
ω_p	0.64	[0.58, 0.69]	0.65	[0.59, 0.71]	0.75	[0.70, 0.80]	0.75	[0.72, 0.79]
ω_w	0.62	[0.55, 0.69]	0.71	[0.66, 0.76]	0.62	[0.56, 0.69]	0.76	[0.72, 0.81]
χ_p	0.48	[0.40, 0.56]	0.49	[0.41, 0.58]	0.43	[0.35, 0.51]	0.50	[0.43, 0.58]
χ_w	0.33	[0.27, 0.39]	0.35	[0.28, 0.41]	0.45	[0.38, 0.53]	0.40	[0.34, 0.47]
ϕ_π	2.01	[1.71, 2.29]	0.39	[0.30, 0.49]	2.19	[1.86, 2.45]	0.57	[0.45, 0.68]
ϕ_y	0.16	[0.09, 0.23]	0.21	[0.17, 0.26]	0.12	[0.04, 0.19]	0.19	[0.15, 0.23]
ρ_r	0.83	[0.78, 0.88]	0.42	[0.31, 0.55]	0.85	[0.81, 0.88]	0.69	[0.61, 0.77]
γ_b^τ	0.04	[0.03, 0.06]	0	–	0.03	[0.02, 0.05]	0	–
γ_b^g	0.05	[0.03, 0.07]	0	–	0.06	[0.04, 0.07]	0	–
γ_y^τ	0.58	[0.38, 0.77]	0.20	[−0.01, 0.40]	0.32	[0.17, 0.48]	0.58	[0.46, 0.68]
γ_y^g	0.31	[−0.05, 0.63]	0.52	[0.32, 0.70]	0.42	[0.08, 0.75]	0.13	[0.08, 0.18]
ρ_g	0.98	[0.97, 1.00]	0.70	[0.47, 0.94]	0.98	[0.97, 0.99]	0.17	[0.04, 0.28]
$\bar{L}$	466.82	[465.61, 468.13]	468.22	[466.67, 469.67]	469.75	[467.54, 471.85]	472.14	[471.00, 473.31]
$\bar{\pi}$	0.90	[0.58, 1.18]	0.62	[0.34, 0.93]	0.65	[0.36, 0.93]	0.66	[0.35, 0.96]
ρ_a	0.35	[0.18, 0.53]	0.38	[0.23, 0.54]	0.45	[0.28, 0.63]	0.41	[0.25, 0.54]
ρ_b	0.31	[0.16, 0.46]	0.33	[0.17, 0.51]	0.33	[0.14, 0.50]	0.50	[0.32, 0.69]
ρ_i	0.43	[0.23, 0.63]	0.48	[0.30, 0.68]	0.72	[0.60, 0.84]	0.73	[0.65, 0.82]
ρ_p	0.73	[0.61, 0.85]	0.90	[0.86, 0.96]	0.88	[0.78, 0.97]	0.92	[0.89, 0.94]
ρ_w	0.34	[0.20, 0.49]	0.50	[0.36, 0.64]	0.27	[0.14, 0.40]	0.19	[0.08, 0.29]
ρ_{tp}	0.45	[0.28, 0.61]	1.00	[0.99, 1.00]	0.11	[0.02, 0.19]	1.00	[0.99, 1.00]
ρ_{em}	0.46	[0.32, 0.61]	0.37	[0.24, 0.52]	0.59	[0.48, 0.68]	0.68	[0.59, 0.78]
$\rho_{e\tau}$	0.73	[0.59, 0.85]	0.82	[0.77, 0.87]	0.95	[0.93, 0.98]	0.84	[0.77, 0.91]
ρ_{eg}	0.40	[0.25, 0.56]	0.71	[0.45, 0.92]	0.26	[0.12, 0.40]	0.98	[0.98, 0.99]
$100\sigma_a$	1.31	[1.13, 1.51]	1.28	[1.07, 1.44]	0.82	[0.72, 0.91]	0.78	[0.69, 0.88]
$100\sigma_b$	6.85	[4.46, 9.17]	11.19	[7.19, 15.36]	11.40	[5.46, 17.48]	8.74	[4.99, 11.85]
$100\sigma_i$	1.23	[0.98, 1.46]	1.21	[0.98, 1.40]	0.46	[0.38, 0.54]	0.51	[0.43, 0.59]
$100\sigma_p$	0.20	[0.16, 0.25]	0.16	[0.12, 0.20]	0.08	[0.06, 0.09]	0.07	[0.05, 0.08]
$100\sigma_w$	0.30	[0.24, 0.37]	0.25	[0.18, 0.32]	0.37	[0.30, 0.43]	0.35	[0.29, 0.40]
$100\sigma_{tp}$	2.98	[2.57, 3.44]	0.14	[0.10, 0.19]	3.00	[2.64, 3.38]	0.22	[0.15, 0.29]
$100\sigma_m$	0.27	[0.22, 0.31]	0.19	[0.16, 0.21]	0.12	[0.11, 0.14]	0.11	[0.10, 0.12]
$100\sigma_\tau^s$	1.31	[1.11, 1.49]	1.42	[1.22, 1.61]	0.60	[0.53, 0.68]	0.59	[0.51, 0.65]
$100\sigma_g$	0.26	[0.22, 0.29]	0.29	[0.24, 0.33]	0.15	[0.13, 0.17]	0.16	[0.14, 0.19]
SP	0.36	[0.26, 0.44]	0.39	[0.29, 0.48]	0.30	[0.21, 0.38]	0.38	[0.29, 0.46]
$\zeta_{sp,b}$	0.05	[0.04, 0.05]	0.05	[0.04, 0.05]	0.04	[0.04, 0.05]	0.05	[0.04, 0.05]
ρ_{σ_ω}	0.94	[0.90, 0.99]	0.92	[0.87, 0.98]	0.99	[0.98, 1.00]	0.95	[0.92, 0.99]
$100\sigma_{\sigma_\omega}$	0.07	[0.06, 0.08]	0.07	[0.06, 0.08]	0.06	[0.05, 0.06]	0.05	[0.05, 0.06]

about the model parameters. Since adding financial frictions to the model and the credit spread to the data does not systematically affect the parameter estimates, we highlight a few observations when comparing the estimates across regimes and subsamples.[11]

First, regarding the nonpolicy parameters, the estimated nominal and real rigidities are quite high to accommodate the stronger-than-usual persistence in the data.[12] For example, while the 90% prior mean of the Calvo parameters (ω_p, ω_w) is 0.6, the post-Volcker sample centers the posterior distributions at 0.75 for both regimes. Moreover, the posterior mean of the habit formation parameter h is estimated to be above 0.9 in all cases, which is significantly higher than its prior mean of 0.5, a result also found by Leeper et al. (2017).

Second, turning to the parameters related to financial frictions, the estimated elasticity of the credit spread with respect to leverage $\zeta_{sp,b}$ is approximately 0.05 in all cases, which is consistent with Christensen and Dib (2008) and Del Negro et al. (2015). Similar to Christiano et al. (2014), the risk shock process turns out to be quite persistent to fit the credit spread data, as the posterior means of ρ_{σ_ω} are all above 0.9.

Finally, moving to the policy parameters, the estimated monetary response to inflation ϕ_π exhibits the typical pattern commonly found in the literature, i.e., higher in the post-Volcker era than in the pre-Volcker era. In addition, the estimated monetary response to output ϕ_y is comparable to Traum and Yang (2011) under regime M but more in line with Leeper et al. (2017) under regime F.

11 A notable exception is the investment adjustment cost parameter s, which is estimated to be smaller than the cases when the model does not feature financial frictions.

12 This result is in line with previous estimates from similar DSGE models, e.g., Smets and Wouters (2007), Justiniano et al. (2011), Traum and Yang (2011), Christiano et al. (2014), and Leeper et al. (2017).

On the other hand, the 90% posterior intervals of $(\gamma_b^\tau, \gamma_b^g)$ all lie within the positive domain under regime M so that both taxes and government spending are systematically used to finance government liabilities. These fiscal instruments also appear to be automatically stabilizing under both regimes and for both subsamples.

3.3. Marginal likelihood

In the Bayesian paradigm, formal regime comparison and selection can be made possible through marginal likelihoods and Bayes factors. This section describes a general numerical strategy for approximating the marginal likelihood implied by high-dimensional DSGE models.

Gelfand and Dey (1994) proposed modifications to the posterior harmonic mean of the likelihood function, thus motivating a simple class of estimators based on the following identity:

$$\frac{1}{p(Y)} = \int \frac{f(\theta)}{p(Y|\theta)p(\theta)} p(\theta|Y) d\theta \tag{3.2}$$

where $f(\theta)$ is a proper density function whose support is contained in that of $p(\theta|Y)$. A popular choice of $f(\theta)$ in economic applications, for example, was proposed by Geweke (1999), who advocated the use of a tail-truncated multivariate Gaussian density tailored to the location and curvature of $p(\theta|Y)$. Given the selection of $f(\theta)$, a simulation-consistent estimator is given by

$$\hat{p}(Y) = \left[\frac{1}{N} \sum_{k=1}^{N} \frac{f(\theta^{(k)})}{p(Y|\theta^{(k)})p(\theta^{(k)})} \right]^{-1} \tag{3.3}$$

where $\{\theta^{(k)}\}_{k=1}^N$ is a sample of N draws from $p(\theta|Y)$. Despite the popularity in empirical macroeconomics, harmonic mean-based estimators raise a number of difficulties when the posterior surface is highly irregular (e.g., nonelliptical and multimodal). One major drawback is its computational instability due to the potentially infinite variance of the estimator; with a sufficiently large N, precarious zeros of $p(\theta|Y)$ in the interior of the parameter space may be picked up by the posterior sampler, thereby exploding the summands in (3.3).

In this paper, we adopt the more reliable method of Chib and Jeliazkov (2001), which is immune to issues with modified harmonic mean estimators. In conjunction with an efficient posterior sampler (e.g., TaRB-MH), this approach provides a general solution to estimating marginal likelihoods in high-dimensional DSGE models. To fix ideas, we start with the basic *marginal likelihood identity* of Chib (1995)

$$p(Y) = \frac{p(Y|\theta)p(\theta)}{p(\theta|Y)} \tag{3.4}$$

which amounts to a rearrangement of Bayes' theorem and holds for all θ. Evaluating the right-hand side of (3.4) at the posterior mode $\hat{\theta}$ and recognizing that the numerator can be easily calculated, the estimation of marginal likelihood reduces to finding an estimate of the posterior ordinate

$$p(\hat{\theta}|Y) = p(\hat{\theta}_1|Y)p(\hat{\theta}_2|\hat{\theta}_1, Y) \cdots p(\hat{\theta}_B|\hat{\theta}_1, \ldots, \hat{\theta}_{B-1}, Y) \tag{3.5}$$

where the parameter space is split into B conveniently specified blocks so that $\hat{\theta} = (\hat{\theta}_1, \ldots, \hat{\theta}_B)$. The decomposition in (3.5) breaks a potentially high-dimensional estimation problem into a set of smaller components that are independently estimable and therefore amenable to parallel computation. Finally, given the estimates $\{\hat{p}(\hat{\theta}_b|\hat{\theta}_1, \ldots, \hat{\theta}_{b-1}, Y)\}_{b=1}^B$ for (3.5), the marginal likelihood on a logarithmic scale can be computed as

$$\ln \hat{p}(Y) = \ln p(Y|\hat{\theta}) + \ln p(\hat{\theta}) - \sum_{b=1}^{B} \ln \hat{p}(\hat{\theta}_b|\hat{\theta}_1, \ldots, \hat{\theta}_{b-1}, Y) \tag{3.6}$$

Practically, we estimate the marginal likelihood from a 6-block parallel procedure based on 10,000 posterior draws for each block. The small numerical standard errors indicate that all marginal likelihoods are tightly estimated.

Several aspects of the implied Bayes factors reported in Tables 4 and 5 are worth highlighting. We focus on discussing the role of financial and fiscal stress in affecting the regime ranking. First, in the absence of financial frictions, the post-Volcker (pre-Volcker) sample overwhelmingly favors regime M (F) in all cases when the deficit-to-debt ratio enters the data (see 'No FF' columns in Table 4). Importantly, all three models identify the switch in monetary policy stance that aligns well with the narrative record of policymakers' beliefs: the passive stance of most of the 1970s under Federal Reserve chairmen Arthur Burns and G. William Miller and the active stance of the early 1980s and beyond under Paul Volcker and Alan Greenspan. With the exception of $\mathcal{M}_1$, which features a persistent component in the tax rule, this result breaks down when the deficit-to-debt ratio is excluded from the data (see 'No FF' columns in Table 5). As in Traum and Yang (2011) and Leeper et al. (2017), both $\mathcal{M}_2$ and $\mathcal{M}_3$ suggest that the data uniformly support regime M over the full sample.

Second, adding financial frictions to the model and credit spread to the data improves the *relative* fit of regime F in all cases, to the extent that it can fundamentally alter the regime ranking (see '+ FF' columns in Tables 4 and 5).[13] Surprisingly or not, regimes

[13] Throughout this section, we treat regime F as the 'null regime' and define its relative fit as the Bayes factor in favor of regime M as opposed to regime F. Thus, an improvement in the relative fit of regime F corresponds to a decrease in the Bayes factor.

Table 4
Log marginal likelihood estimates, with fiscal stress.

Model	1962:Q1–1979:Q2				1984:Q1–2007:Q4			
	No FF	ln BF	+ FF	ln BF	No FF	ln BF	+ FF	ln BF
$\mathcal{M}_1$: **Benchmark**								
Regime-M	−805.17		−725.63		−714.03		−589.67	
	(0.06)	−12.06*	(0.06)	−20.59*	(0.07)	12.49*	(0.10)	4.34
Regime-F	−793.11		−705.04		−726.52		−594.01	
	(0.09)		(0.12)		(0.08)		(0.08)	
$\mathcal{M}_2$: **LTW + distorting tax**								
Regime-M	−947.69		−868.94		−925.47		−801.89	
	(0.08)	−15.66*	(0.08)	−21.92*	(0.07)	15.54*	(0.10)	5.72*
Regime-F	−932.03		−847.02		−941.01		−807.61	
	(0.08)		(0.09)		(0.15)		(0.09)	
$\mathcal{M}_3$: **LTW + lump-sum tax**								
Regime-M	−967.29		−886.59		−933.01		−805.75	
	(0.08)	−14.85*	(0.07)	−25.49*	(0.06)	26.68*	(0.08)	17.47*
Regime-F	−952.44		−861.10		−959.69		−823.22	
	(0.08)		(0.06)		(0.08)		(0.09)	

Notes: Marginal likelihood estimates with numerical standard errors in parentheses and Bayes factors (BF) are reported in logarithm scale for $\mathcal{M}_1$–$\mathcal{M}_3$ under each regime both with and without financial frictions (FF). Asterisk (*) signifies decisive evidence in favor of the regime with superior fit, corresponding to a log Bayes factor greater than 4.6 in absolute value based on Jeffreys's (1961) criterion. LTW stands for Leeper et al. (2017).

Table 5
Log marginal likelihood estimates, without fiscal stress.

Model	1962:Q1–1979:Q2				1984:Q1–2007:Q4			
	No FF	ln BF	+ FF	ln BF	No FF	ln BF	+ FF	ln BF
$\mathcal{M}_1$: **Benchmark**								
Regime-M	−700.05		−621.62		−660.07		−535.94	
	(0.06)	−14.92*	(0.06)	−18.80*	(0.08)	12.67*	(0.11)	0.82
Regime-F	−685.13		−602.82		−672.74		−536.76	
	(0.08)		(0.11)		(0.06)		(0.07)	
$\mathcal{M}_2$: **LTW + distorting tax**								
Regime-M	−815.49		−735.32		−861.51		−732.56	
	(0.08)	11.71*	(0.08)	6.96*	(0.09)	24.92*	(0.09)	21.86*
Regime-F	−827.20		−742.28		−886.43		−754.42	
	(0.10)		(0.08)		(0.14)		(0.14)	
$\mathcal{M}_3$: **LTW + lump-sum tax**								
Regime-M	−839.68		−756.70		−867.33		−735.97	
	(0.10)	8.95*	(0.08)	0.88	(0.09)	37.92*	(0.11)	34.69*
Regime-F	−848.63		−757.58		−905.25		−770.66	
	(0.10)		(0.08)		(0.13)		(0.12)	

Notes: See Table 4.

M and F become 'nearly' observationally equivalent in $\mathcal{M}_1$ over the post-Volcker era, for which the tenability of regime M has rarely been challenged in the literature. This observational equivalence result also holds for $\mathcal{M}_3$ over the pre-Volcker era without including the deficit-to-debt ratio in the data. Taken together, these findings echo the theoretical demonstration of Leeper and Walker (2013) that the two policy regimes can provide equally plausible interpretations of the data. Consequently, policymakers should routinely scrutinize alternative monetary–fiscal policy specifications in their policymaking process.

3.4. Inspecting the mechanism

Given the empirical relevance of financial frictions for regime identification, a natural question is, then, what are the main features in the data that regime F can explain but regime M cannot when an otherwise standard new Keynesian DSGE model is extended to incorporate financial frictions. To gain some preliminary insights, it is informative to examine the comovements between the credit spread, which is newly appended to the data, and other existing observables for which the two policy regimes produce qualitatively different predictions. For example, Fig. 2 compares the cross-correlations between the lagged credit spread and inflation that are implied by the data with their model-implied counterparts. The pre-Volcker sample (Panel A) exhibits prolonged positive covariations for at least 8 years before it turns negative. According to the posterior means of model parameters, this feature only emerges in regime F, although its model-implied correlations appear much weaker and relatively short-lived compared to the data. With a one-year lag, the post-Volcker sample (Panel B) also displays a similar pattern of covariations that seems to be better captured by regime F.

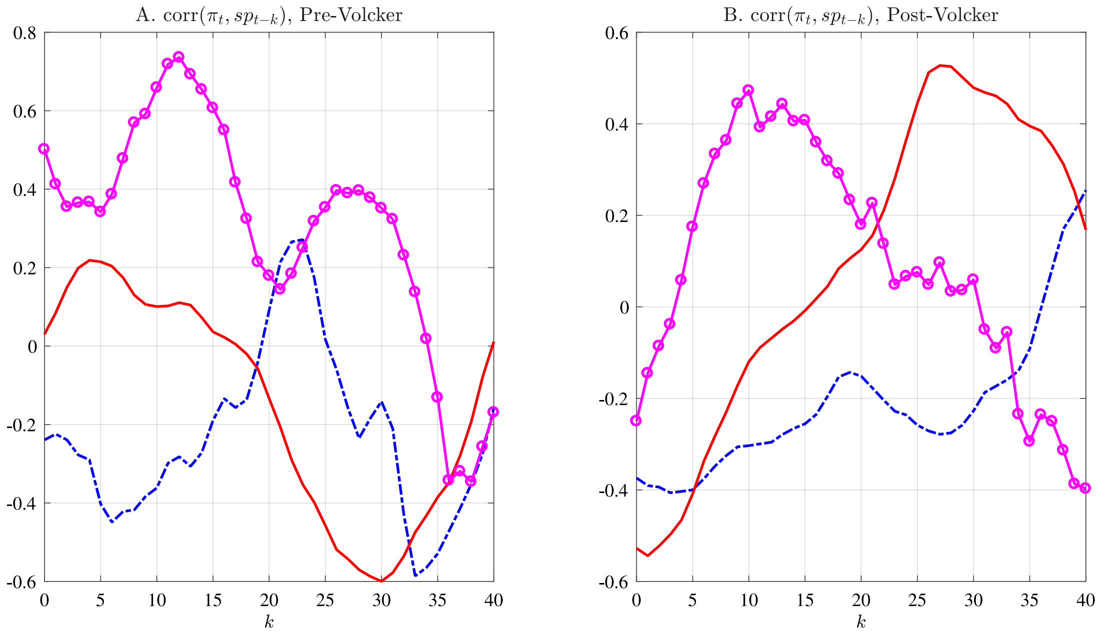

Fig. 2. Cross-correlation between inflation (π) and credit spread (sp), $\mathcal{M}_1$ with financial frictions. Notes: Both panels compare the cross-correlation functions computed from the actual data (purple line with circle) and the simulated data of $\mathcal{M}_1$ evaluated with the posterior means under regime M (blue dashed line) and regime F (red solid line). The simulated data have the same sample size as the actual data. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

Of course, since the macroeconomy amounts to a sophisticated system disturbed by various exogenous forces, there can be multiple factors that give rise to the cross-regime divergence shown in Fig. 2. To pinpoint these factors, we launch a series of counterfactual exercises by simulating artificial data from the model with one structural shock turned off at a time and then calculate the model-implied cross-correlations based on the simulated data. We identify several shocks, including risk, monetary policy, and tax shocks, as the most important factors because the cross-correlograms between credit spread and inflation change drastically when each of them is muted. In what follows, we provide a close look at the impulse responses of $\mathcal{M}_1$ with financial frictions to these exogenous shocks.[14]

3.4.1. Risk shock

Fig. 3 displays the dynamic responses of key model variables to an adverse risk shock. Since banks charge a higher interest rate on loans in response to a rise in risk, the credit extended to entrepreneurs falls. With fewer financial resources, entrepreneurs purchase less physical capital from the capital goods producer, who in turn acquires fewer investment goods. Under both regimes, falling investment leads to a drop in output; it also pushes down the price of physical capital and hence the net worth of entrepreneurs, amplifying the impacts of credit market tensions on the macroeconomy through the standard financial accelerator.

Turning to the policy reactions to the economic contraction, the monetary authority responds by more aggressively cutting the policy rate in regime M. With sticky prices, a lower nominal interest rate reduces the real interest rate. As a result, consumption in regime M rises by more than that in regime F, tempering the short-run drop in output. The automatic stabilizing fiscal instruments, on the other hand, respond by cutting taxes and increasing expenditures financed by nominal bond sales. Together with lower output and higher bond price due to lower nominal rate, this fiscal expansion raises the market value of debt as a share of output in both regimes. Because higher deficits do not trigger expectations of sufficiently high surpluses to stabilize debt in regime F, households feel wealthier and demand more consumption goods, which bids up the price level. The resulting higher inflation and lower bond price jointly ensure that the market value of debt is aligned with the expected present value of surpluses. Such a wealth effect, if any, will almost be neutralized by the passive fiscal policy in regime M so that the decline in economic activity results in a decline in the marginal cost of production and hence inflation.

In sum, the two policy regimes produce strikingly different inflation dynamics following a credit crunch, which, to a large extent, explains the outperformance of regime F in fitting the cross-correlations displayed in Fig. 2. In particular, contrary to regime M, which underlies the standard new Keynesian analysis such as Christiano et al. (2014), elevated credit risk can bring forth fiscal inflation through the debt revaluation channel that regime F emphasizes. The resulting comovements between the lagged credit spread and inflation concur with the overall positive covariations observed in the data.

[14] To conserve space, impulse responses to other shocks of the model are not reported here and will be available from the authors upon request.

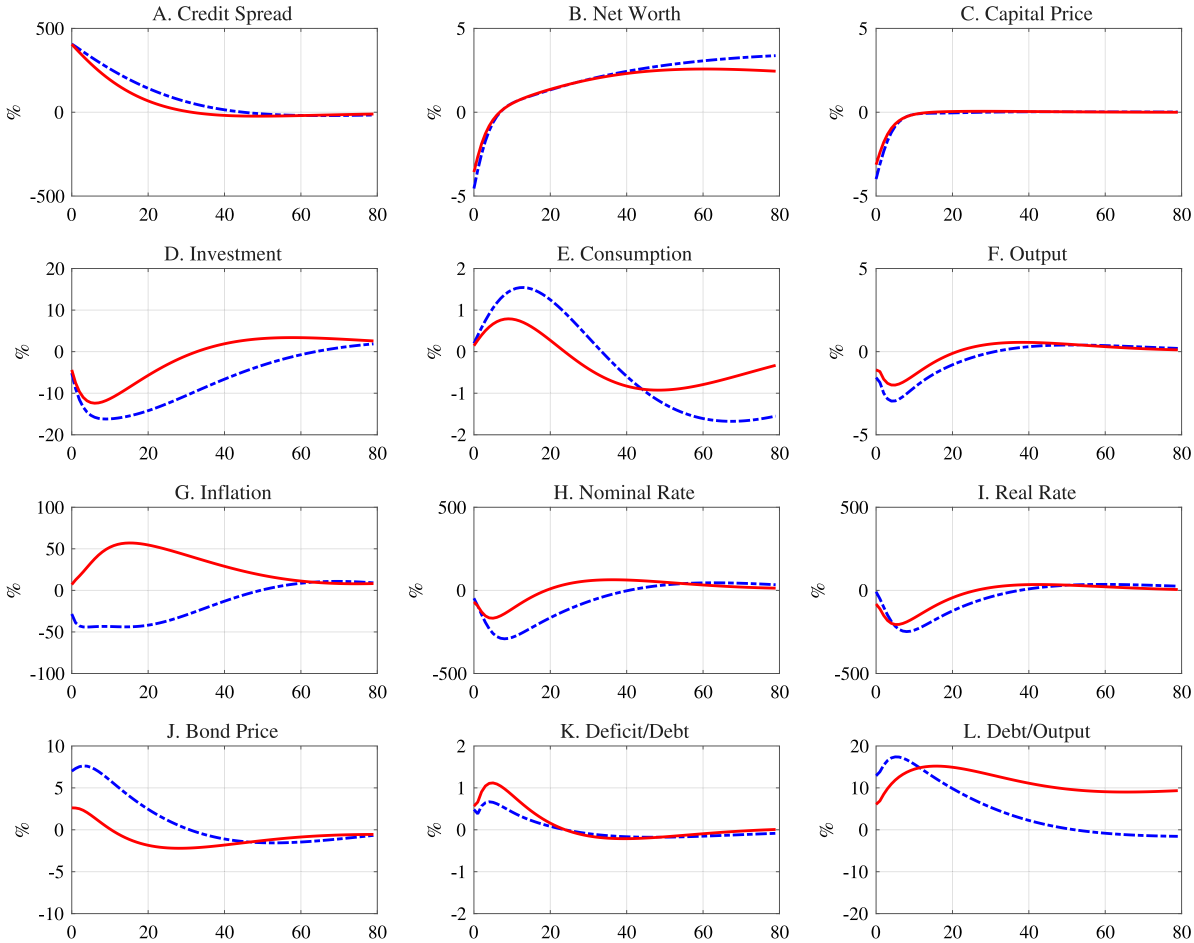

Fig. 3. Dynamic responses to a one percent positive risk shock, $\mathcal{M}_1$ with financial frictions. Notes: Responses are evaluated at the posterior mean parameters of regime M (blue dashed line) and regime F (red solid line) based on the full sample. Credit spread, inflation, nominal and real rates are converted into annualized basis points; the remaining variables are in percentage deviations from the steady state. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

3.4.2. Interest rate shock

Next, Fig. 4 displays the dynamic responses of key model variables to a contractionary monetary policy shock. Given sticky prices, a higher nominal interest rate translates into a higher real interest rate that makes consumption today more costly relative to tomorrow. Meanwhile, the higher real rate dampens the demand for capital, which explains the decrease in investment and hence the price of capital. As a result, consumption, investment, and output fall immediately under both regimes. The lower capital price also reduces the net worth of entrepreneurs, thereby deteriorating their balance sheet conditions and pushing up the credit spread in the very short run. On the other hand, while inflation falls on impact under regime M as conventional monetary theory predicts, it rises immediately and persistently through the debt revaluation channel of regime F, as in the case of risk shock. As a result, the overall positive covariations between lagged credit spread and inflation that arise from regime F align with the observed patterns in the data.

3.4.3. Tax shock

Finally, Fig. 5 displays the impulse responses of key model variables to a deficit-financed tax cut. Due to the positive wealth effect, this fiscal expansion brings about an immediate and persistent rise in inflation under both regimes. In particular, inflation rises under regime M because the present value of expected higher future taxes does not exactly neutralize today's asset increase for the benchmark model. Such a wealth effect becomes even stronger so that inflation almost doubles under regime F since the private agent anticipates no future tax adjustment to finance the rising government debt. On the other hand, higher inflation lowers the real interest rate, and it does so more strongly under regime F, which in turn stimulates the demand for capital goods. Meanwhile, the supply of capital goods falls as more resources are extracted by the government. Because the effects of higher demand (lower supply) dominate in the capital goods market under regime F (M), installed capital and investment increase (decrease) with a rise in the price of capital. A higher capital price raises the net worth of entrepreneurs, thereby improving their balance sheet conditions and

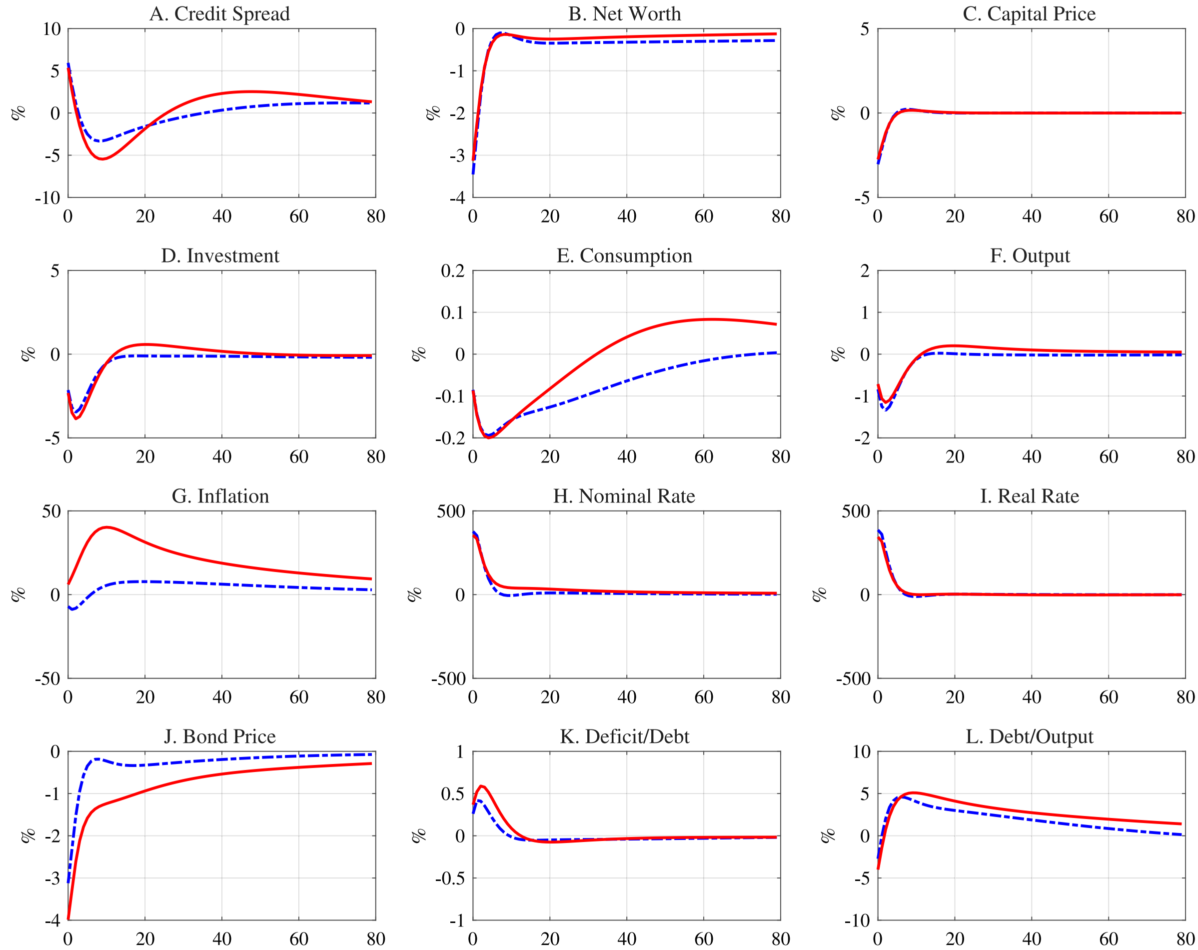

Fig. 4. Dynamic responses to a one percent positive interest rate shock, $\mathcal{M}_1$ with financial frictions. Notes: See Fig. 3.

pushing down the credit spread on impact. However, movements in credit spread are also linked endogenously to installed capital via entrepreneurs' leverage as (2.2) depicts. With higher (lower) installed capital, leverage, and hence credit spread, rise (fall) under regime F (M). Taken together, the overall positive covariations between the lagged credit spread and inflation that regime F yields are in line with the observed comovements in the data.

4. Confronting policy regime uncertainty

Our foregoing analysis operates under an implicit assumption that the underlying model space is complete—one of the models under regime M or F is correctly specified. As a consequence, the full posterior weight will be assigned to whichever model lies closest (in terms of the Kullback–Leibler divergence) to the true data generating process as the sample size $T \to \infty$ [see, e.g., Geweke and Amisano, 2011]. However, more realistically, a prudent policymaker may wish to treat each model as misspecified in some aspects of reality and therefore base her policy thinking beyond the implications of any single policy regime. Indeed, the marginal likelihood estimates in Tables 4 and 5 reveal that even similar model specifications or datasets can produce widely different rankings of the policy regime. This gives rise to the natural question of regime selection or composition for policymakers.

To confront policy regime uncertainty, this section assumes an incomplete model space by dynamically combining both policy regimes in a linear prediction pool model. In particular, our approach integrates those of Waggoner and Zha (2012) and Del Negro et al. (2016) in two aspects. First, we employ an autoregressive process that induces smooth rather than drastic changes in the regime weights. Second, we allow for the joint estimation of model parameters under both regimes as well as the regime weights. To fix the idea, let θ_R collect the model parameters under regime $R \in \{M, F\}$ and $Y_{1:T}$ collect the sample observations y_t for periods $t = 1, \ldots, T$. Suppose that the policymaker dynamically pools the predictive densities generated by each policy regime according to the following mixture distribution:

$$p(y_t|\theta_M, \theta_F, \lambda_t, Y_{1:t-1}) = \lambda_t p(y_t|\theta_M, Y_{1:t-1}) + (1 - \lambda_t)p(y_t|\theta_F, Y_{1:t-1}) \quad (4.1)$$

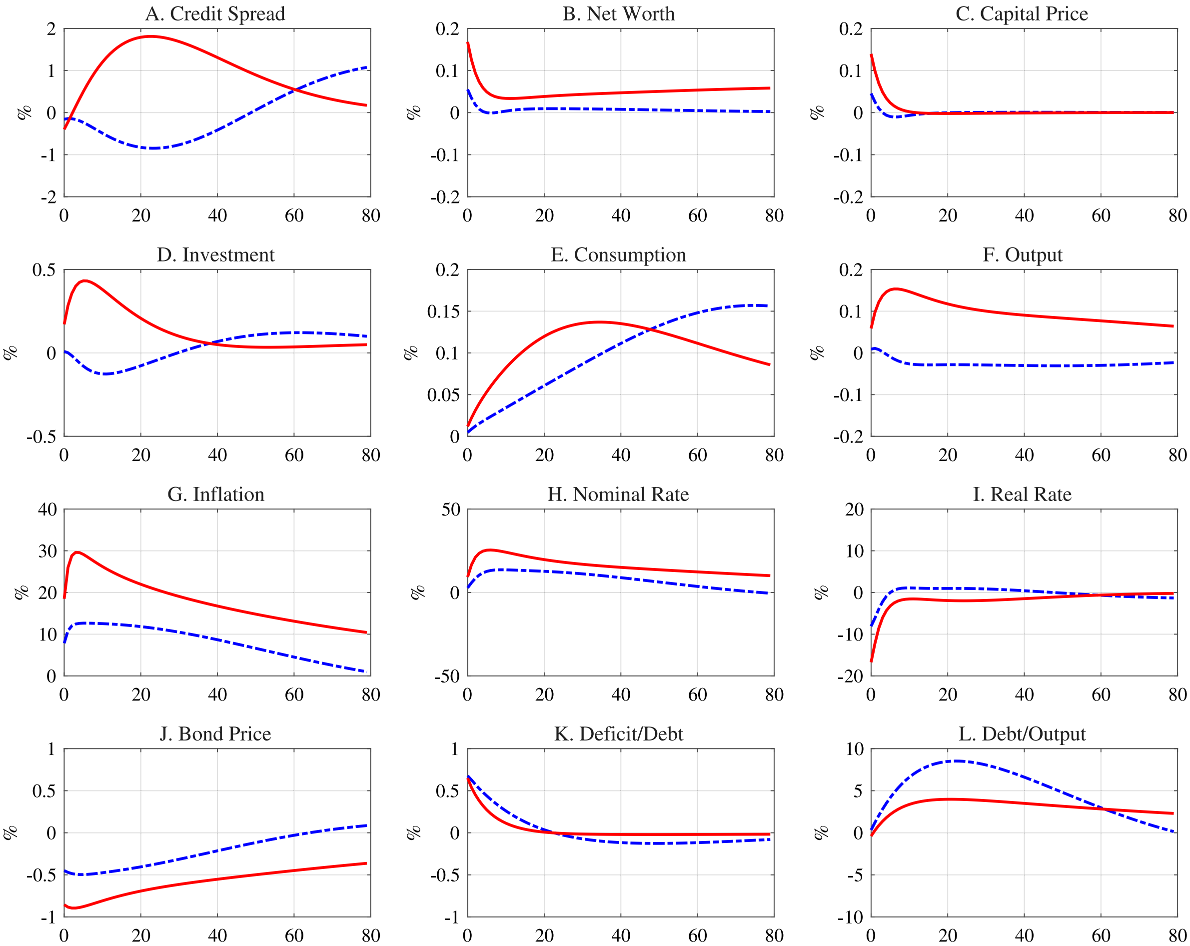

Fig. 5. Dynamic responses to a one percent negative tax shock, $\mathcal{M}_1$ with financial frictions. Notes: See Fig. 3.

where $\lambda_t = \Phi(x_t)$ measures the model weight of regime M and $\Phi(\cdot)$ denotes the cumulative distribution function of a standard normal random variable. In addition, x_t is an autoregressive latent factor

$$x_t = (1-\rho)\mu + \rho x_{t-1} + \sqrt{1-\rho^2}v_t, \qquad v_t \sim \mathbb{N}(0,\sigma^2), \qquad x_0 \sim \mathbb{N}(\mu,\sigma^2) \tag{4.2}$$

where the following priors for the hyperparameters are used: $\mu \sim \mathbb{N}(0, 0.5^2)$, $\rho \sim \mathbb{U}(0,1)$, and $\sigma \sim \mathbb{IG}\text{-}1(4.175, 0.807)$. In the Online Appendix, we develop a posterior sampling algorithm for estimating this dynamic prediction pool of both policy regimes.[15]

Fig. 6 plots the historical weight series for regime M estimated over the full sample. All series are evaluated at their posterior means under $\mathcal{M}_1$ (black solid line), $\mathcal{M}_2$ (blue dashed line), and $\mathcal{M}_3$ (red dash–dot line) with financial frictions.[16] It displays prima facie evidence of time variations in the relative importance of each regime. Unlike the clear-cut regime ranking based on marginal likelihood comparison, both regimes receive nondegenerate pooling weights throughout the entire sample period. Fig. 6 also highlights an interesting pattern of transitions across regimes. Despite the differences in model specification and dataset, all three series exhibit pronounced cyclical fluctuations, with marked decreases in the relevance of regime M (or equivalently, sharp increases in the importance of regime F) accompanying the recessions as designated by the National Bureau of Economic Research. While our full sample does not span the 2007–2009 Great Recession period, the estimated pooling weights of regime M already show a declining trend since the early 2000s. On average, regime M appears to be more relevant during the Great Moderation period.

[15] See also Tan (2019) for a frequency-domain implementation of this approach.

[16] To conserve space, we only plot the regime weights when the deficit-to-debt ratio is excluded from the data. The results are largely unchanged when this observable is included in the data.

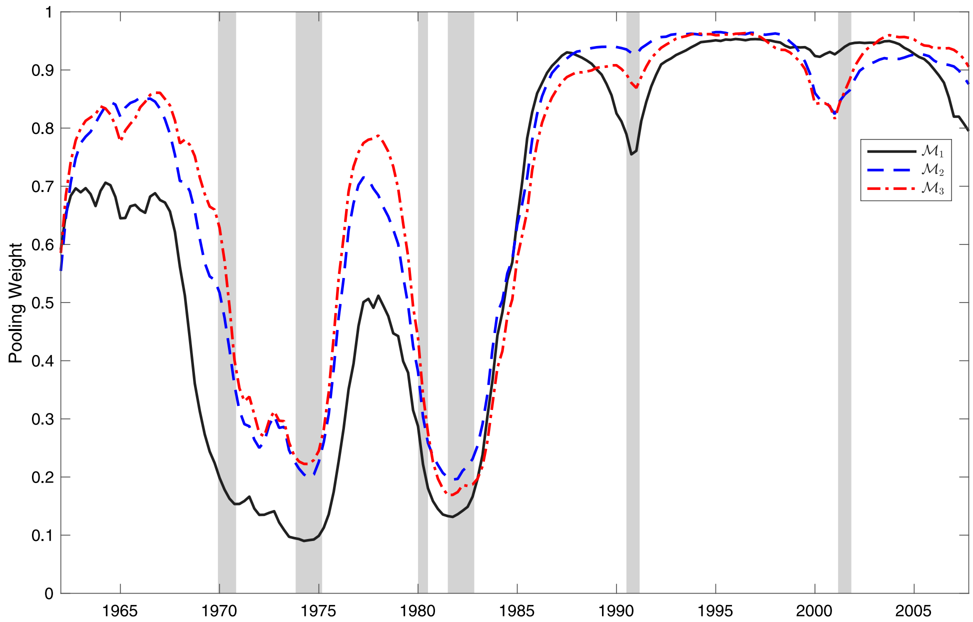

Fig. 6. Estimated pooling weights for regime M. Notes: All weights are evaluated at their posterior means under $\mathcal{M}_1$ (black solid line), $\mathcal{M}_2$ (blue dashed line), and $\mathcal{M}_3$ (red dash–dot line) with financial frictions. Shaded bars indicate recessions as designated by the National Bureau of Economic Research. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

5. Concluding remarks

This article studies monetary and fiscal policy interactions in a financially constrained environment. We estimate various versions of a standard medium-sized DSGE model augmented with fiscal details and credit market imperfections and find that adding financial frictions to the model and financial variables to the dataset plays an important role in improving the relative fit of regime F that embodies the fiscal theory. The improvement stems largely from regime F's implications for the comovement patterns between credit spread and inflation, which find empirical support in the data. To confront policy regime uncertainty, we advocate the use of linear prediction pools that dynamically combine both policy regimes. Alternatively, further identifying restrictions from elsewhere is still necessary to break the near observational equivalence of the two policy regimes found in the present study.

CRediT authorship contribution statement

Bing Li: Conceptualization, Methodology, Investigation, Writing – original draft, Funding acquisition. **Pei Pei:** Validation, Writing – original draft, Project administration. **Fei Tan:** Conceptualization, Software, Formal analysis, Writing – review & editing, Funding acquisition.

Appendix A. Supplementary material

Supplementary material related to this article can be found online at https://doi.org/10.1016/j.jmacro.2021.103353.

References

Bernanke, B.S., Gertler, M., Gilchrist, S., 1999. The financial accelerator in a quantitative business cycle framework. In: Handbook of Macroeconomics, vol. 1, Elsevier, pp. 1341–1393.
Bhattarai, S., Lee, J.W., Park, W.Y., 2012. Monetary-fiscal policy interactions and indeterminacy in postwar US data. Am. Econ. Rev. Pap. Proc. 102 (3), 173–178.
Bhattarai, S., Lee, J.W., Park, W.Y., 2016. Policy regimes, policy shifts, and U.S. business cycles. Rev. Econ. Stat. 98 (5), 968–983.
Bianchi, F., 2012. Evolving monetary/fiscal policy mix in the United States. The Am. Econ. Rev. Pap. Proc. 102 (3), 167–172.
Bianchi, F., 2013. Regime switches, agents' beliefs, and post-world war II U.S. macroeconomic dynamics. Rev. Econom. Stud. 80 (2), 463–490.
Bianchi, F., Ilut, C., 2017. Monetary/fiscal policy mix and agent's beliefs. Rev. Econ. Dyn. 26, 113–139.
Chen, X., Leeper, E., Leith, C., 2019. US monetary and fiscal policies—Conflict or cooperation? Manuscript.
Chib, S., 1995. Marginal likelihood from the gibbs output. J. Amer. Statist. Assoc. 90, 1313–1321.
Chib, S., Jeliazkov, I., 2001. Marginal likelihood from the Metropolis-Hastings output. J. Amer. Statist. Assoc. 96 (453), 270–281.
Chib, S., Ramamurthy, S., 2010. Tailored randomized block MCMC methods with application to DSGE models. J. Econometrics 155 (1), 19–38.
Chib, S., Shin, M., Tan, F., 2021. DSGE-SVt: An Econometric Toolkit for High-Dimensional DSGE Models with SV and *t* Errors. Manuscript.
Chib, S., Zeng, X., 2019. Which factors are risk factors in asset pricing? A model scan framework. J. Bus. Econom. Statist. 1–28.

Christensen, I., Dib, A., 2008. The financial accelerator in an estimated new keynesian model. Rev. Econ. Dyn. (ISSN: 1094-2025) 11, 155–178. http://dx.doi.org/10.1016/j.red.2007.04.006.
Christiano, L.J., Eichenbaum, M., Evans, C.L., 2005. Nominal rigidities and the dynamic effects of a shock to monetary policy. J. Polit. Econ. 113 (1), 1–45.
Christiano, L., Motto, R., Rostagno, M., 2003. The great depression and the friedman-schwartz hypothesis. J. Money Credit Bank. 35 (6), 1119–1197.
Christiano, L.J., Motto, R., Rostagno, M., 2014. Risk shocks. Amer. Econ. Rev. 104 (1), 27–65.
Clarida, R., Galí, J., Gertler, M., 2000. Monetary policy rules and macroeconomic stability: Evidence and some theory. Q. J. Econ. 115 (1), 147–180.
Cochrane, J.H., 1999. A frictionless view of U.S. inflation. In: Bernanke, B.S., Rotemberg, J.J. (Eds.), NBER Macroeconomics Annual 1998, Vol. 14. MIT Press, Cambridge, MA, pp. 323–384.
Cui, W., 2016. Monetary-fiscal interactions with endogenous liquidity frictions. Eur. Econ. Rev. 87, 1–25.
Davig, T., Leeper, E.M., 2006. Fluctuating macro policies and the fiscal theory. In: Acemoglu, D., Rogoff, K., Woodford, M. (Eds.), NBER Macroeconomics Annual 2006, Vol. 21. MIT Press, Cambridge, pp. 247–298.
Del Negro, M., Giannoni, M.P., Schorfheide, F., 2015. Inflation in the great recession and new keynesian models. Am. Econ. J. Macroecon. 7 (1), 168–196.
Del Negro, M., Hasegawa, R.B., Schorfheide, F., 2016. Dynamic prediction pools: An investigation of financial frictions and forecasting performance. J. Econometrics 192 (2), 391–405, Innovations in Multiple Time Series Analysis.
Gelfand, A.E., Dey, D.K., 1994. Bayesian model choice: Asymptotics and exact calculations. J. R. Stat. Soc. Ser. B Stat. Methodol. 56 (3), 501–514.
Geweke, J., 1999. Using simulation methods for Bayesian econometric models: Inference, development, and communication. Econometric Rev. 18, 1–73.
Geweke, J., Amisano, G., 2011. Optimal prediction pools. J. Econometrics 164 (1), 130–141, Annals Issue on Forecasting.
Gomes, P., Seoane, H.D., 2018. Made in Europe: Monetary-fiscal policy mix with financial frictions. Manuscript, Universidad Carlos III de Madrid.
Jeffreys, J., 1961. Theory of Probability. Oxford University Press.
Justiniano, A., Primiceri, G.E., Tambalotti, A., 2011. Investment shocks and the relative price of investment. Rev. Econ. Dyn. (ISSN: 1094-2025) 14, 102–121. http://dx.doi.org/10.1016/j.red.2010.08.004.
Kliem, M., Kriwoluzky, A., Sarferaz, S., 2016a. Monetary-fiscal policy interaction and fiscal inflation: A tale of three countries. Eur. Econ. Rev. 88, 158–184.
Kliem, M., Kriwoluzky, A., Sarferaz, S., 2016b. On the low-frequency relationship between public deficits and inflation. J. Appl. Econometrics 31 (3), 566–583.
Leeper, E.M., 1991. Equilibria under 'active' and 'passive' monetary and fiscal policies. J. Monetary Econ. 27 (1), 129–147.
Leeper, E.M., Li, B., 2017. Surplus-debt regressions. Econom. Lett. 151, 10–15.
Leeper, E.M., Traum, N., Walker, T.B., 2017. Clearing up the fiscal multiplier morass. Amer. Econ. Rev. 107 (8), 2409–2454.
Leeper, E., Walker, T., 2013. Perceptions and misperceptions of fiscal inflation. In: Fiscal Policy After the Financial Crisis. National Bureau of Economic Research, Inc, pp. 255–299.
Li, E.X.N., Zha, T.A., Zhang, J., Zhou, H., 2018. Active monetary or fiscal policy and stock-bond correlation. Manuscript.
Lubik, T.A., Schorfheide, F., 2004. Testing for indeterminacy: An application to U.S. monetary policy. Amer. Econ. Rev. 94 (1), 190–217.
Sims, C.A., 2002. Solving linear rational expectations models. Comput. Econ. 20 (1), 1–20.
Sims, C.A., 2011. Stepping on a rake: The role of fiscal policy in the inflation of the 1970's. Eur. Econ. Rev. 55 (1), 48–56.
Sims, C.A., 2012. Modeling the influence of fiscal policy on inflation. Manuscript, Princeton University.
Sims, C.A., 2013. Paper money. Amer. Econ. Rev. 103 (2), 563–584.
Smets, F., Wouters, R., 2007. Shocks and frictions in US business cycles: A Bayesian DSGE approach. Amer. Econ. Rev. 97 (3), 586–606.
Tan, F., 2019. A frequency-domain approach to dynamic macroeconomic models. Macroeconomic Dynamics 1–31.
Traum, N., Yang, S.-C.S., 2011. Monetary and fiscal policy interactions in the post-war U.S.. Eur. Econ. Rev. 55 (1), 140–164.
Waggoner, D.F., Zha, T., 2012. Confronting model misspecification in macroeconomics. J. Econometrics 171 (2), 167–184.
Woodford, M., 1995. Price-level determinacy without control of a monetary aggregate. Carneige-Rochester Conf. Ser. Public Policy 43, 1–46.
Woodford, M., 2001. Fiscal requirements for price stability. J. Money Credit Bank. 33 (3), 669–728.
Xu, L., Serletis, A., 2017. Financial frictions and the fiscal theory of price level determination. Open Econ. Rev. 28 (2), 251–272.