From Conventional to Unconventional Monetary Policy: Is the Taylor Rule an Adequate Representation in Macro Models?

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Abstract

Probably not. A Taylor Rule remains the consensus in macro models despite unconventional monetary policies (UMP) and zero funds rate in 2009-2015. Using three benchmark models - a vector autoregression, a small-scale New Keynesian model, and a medium-scale DSGE model - we test for structural breaks between the Great Moderation and post-crisis periods. Structural breaks occurred at 2007:Q3 in the Taylor Rule (less "hawkish" than the Great Moderation) and many non-policy structural parameters, which exhibit significant time-variation. We examine whether these breaks reflect omitted UMP effects by comparing results using the federal funds rate (with a zero-lower bound constraint) and Wu-Xia shadow rate. Surprisingly, model estimation and projection is qualitatively similar using either rate, but dynamic properties are inconsistent across models. This suggests either that structural breaks are not due to UMP or that the shadow rate is an ineffective UMP proxy. Explicit specification of UMP structure in macro models is likely necessary.

Keywords: Taylor Rule, structural break, macroeconomic models, unconventional monetary policy

JEL Classification: E43 , E52 , E58 , E12 , E13 , E65 , E61 , C50 , C32

1 Introduction

More than a decade after the Global Financial Crisis (GFC) and the adoption of unconventional monetary policies (UMP), John Taylor's (1993) interest rate rule remains the dominant framework for modeling monetary policy in macroeconomics. It features prominently in the benchmark New Keynesian DSGE model (Vines and Wills, 2018) and in macroeconomic textbooks at all levels. Even recent theoretical innovations continue to rely on Taylor Rule-like representations, highlighting its enduring influence despite dramatic changes in monetary policy regimes over the past 15 years.

The Federal Reserve still relies on a prototypical Taylor Rule in its primary macro model, FRB/US:

$$i_t = \rho_i i_{t-1} + (1 - \rho_i)[r + \pi_t + \phi_\pi(\pi_t - \pi^*) + \phi_y y_t] + \epsilon_t \tag{1}$$

where i_t is the nominal federal funds interest rate, r is the "natural" (equilibrium) real rate, π_t is inflation, π^* is target inflation, and y_t is the output gap (Brayton et al, 2014).

Despite its widespread adoption, the sufficiency of the Taylor Rule as a representation of monetary policy has come under scrutiny. The GFC and subsequent period of zero interest rates prompted the Federal Reserve to deploy a number of historically unconventional monetary policy (UMP) tools that greatly expanded the conventional interest rate targeting framework captured by the Taylor Rule.

First, the Fed began paying banks interest on reserves (IOR) in 2008 to improve its control over the monetary base. Second, the Fed implemented large-scale asset purchases (LSAP) of Treasury securities and mortgage-backed securities totaling over \$3 trillion between 2008-2014. LSAPs, also referred to as Quantitative Easing (QE), target long-term yields rather than operating solely through short-term rates and money demand. Third, forward guidance emerged as a distinct policy tool, with explicit communication about future policy intentions itself becoming a policy tool. Fourth, the Fed expanded its toolkit with new lending facilities and direct credit market interventions that bypass traditional monetary transmission channels entirely.

These UMP tools operate through fundamentally different transmission mechanisms compared to conventional policy. While the Taylor Rule assumes monetary policy works through a single policy instrument (short-term interest rate) that affects broader economic conditions indirectly and imprecisely, UMP added additional policy instruments (reserves, long-term interest rates, lending to banks, and bank financial stability) that influences broader economic conditions more directly effectively. However, the addition of UMP did not include additional specific targets, objectives, or rules. For example, the Fed's balance sheet expanded from roughly $900 billion precrisis to over $4 trillion by 2014—a transformation that is not likely captured by single short-term interest rate rule alone.

These UMP additions raise a fundamental question about whether the Taylor Rule—designed for conventional policy environments where the central bank's primary tool is a short-term interest rate—can comprehensively characterize monetary policy during and after the crisis. This question reflects two distinct challenges in representing the "art of central banking" through a simple equation. First, even under conventional monetary policy, a single interest rate rule offers only a partial view given that central banks employ various instruments beyond just the policy rate, such as reserve requirements, discount window operations, and open market operations. Second, the introduction of UMP has expanded the policy toolkit beyond the short-term interest rate, adding new instruments such as quantitative easing and forward guidance that operate through different transmission channels. Bernanke (2020) argues that "old methods won't do" for modern monetary policy, calling for updated tools and frameworks that can accommodate this expanded policy space. The omission of these UMP tools in standard macroeconomic models may lead to structural breaks in estimated policy rules that reflect model misspecification rather than genuine changes in central bank behavior.

This paper contributes to the debate by evaluating whether the Taylor Rule remains a useful representation of monetary policy in macroeconomic models after the GFC. Much of the existing structural break literature has primarily documented parameter instability during earlier monetary policy regimes but has not systematically examined the post-2007 period when UMP fundamentally altered the policy landscape (see Table 1). Studies like Clarida et al (2000), Canova (2009), and Ilbas (2012) focused on parameter changes during the Volcker disinflation and Great Moderation, finding that the Fed became more responsive to inflation. More recent work has documented structural breaks after 2007 (Nikolsko-Rzhevskyy et al, 2019) but has not systematically tested whether these breaks are attributable to UMP omission—a critical distinction for model specification and policy analysis.

Our analysis examines whether omission of UMPs drives the widespread structural breaks documented after 2007, using multiple model classes and shadow rate controls to isolate the effects of unconventional policies. We test for structural breaks in three benchmark models—a vector autoregression (VAR), a New Keynesian (NK) model, and a dynamic stochastic general equilibrium (DSGE) model—between the Great Moderation and post-crisis periods. To approximate the effects of UMP, we incorporate the shadow federal funds rate proposed by Wu and Xia (2016), which adjusts the Taylor Rule to account for unconventional policies indirectly. This approach allows us to assess whether structural breaks observed with the actual funds rate persist when UMP is considered.

We also evaluate whether controlling for the zero lower bound (ZLB) constraint is sufficient to address the observed structural breaks. Our approach follows Cuba-Borda et al (2019) and Giovannini et al (2021) in the NK and DSGE models, which impose

ZLB constraints during periods of low interest rates. In addition, we control for the ZLB in the VAR model using the SVAR with occasionally binding constraints method developed by Aruoba et al (2022). These strategies accommodate non-linearities at the ZLB while allowing for tractability and comparability across frameworks.

The results indicate that structural breaks occur not only in Taylor Rule coefficients but also in non-policy parameters, suggesting broader changes in the macroeconomic environment. Models estimated with shadow rates show evidence of significant breaks but capture some effects of UMP more effectively than those using the federal funds rate alone. While the shadow rate reduces some of the discrepancies, it does not fully eliminate structural breaks, implying that unconventional monetary policies may operate through channels not fully accounted for by standard Taylor Rules. In addition, results controlling for the ZLB produce breaks similar in magnitude to those observed with the shadow rate. This raises the question of whether non-linearities at the ZLB are sufficient to explain the observed dynamics, or whether more explicit modeling of UMP tools is necessary.

The estimated coefficients may fail to accurately reflect policymakers' true preferences regarding inflation control versus output stabilization. Additionally, central banks may pursue objectives beyond the traditional inflation-output trade-off, including financial stability, exchange rate management, or distributional concerns that are absent from standard Taylor Rules. Most fundamentally, the proliferation of policy instruments, from asset purchases to forward guidance, highlight the limitations of relying on conventional Taylor Rules even when adjusted for UMP proxies and point to the need for richer structural frameworks to address modern policy complexities.

The paper proceeds as follows. Section 2 reviews the literature on structural breaks in monetary policy, including approaches to modeling UMP and the ZLB. Section 3 outlines the models used in the analysis, and Section 4 describes the estimation methods. Section 5 presents results, while Section 6 provides additional diagnostics and sensitivity checks. Section 7 discusses broader implications, and Section 8 concludes with directions for future research.

2 Previous Literature

Table 1 lists the main papers reporting evidence on structural breaks in the Taylor Rule[1]. The structure of the Rule has stayed largely the same as the original specification except for the addition of persistence (i_{t-1}), allowance for output dynamics (growth rate or gap change), and variation in lags or other practical features, as shown in Coibion and Gorodnichenko (2012).[2] The literature contains a variety of different modeling and econometric methods used to estimate breaks in the Taylor Rule coefficients. However, the results tend to be broadly consistent across papers.

Regime-switching models and break point tests, like those in Estrella and Fuhrer (2003) and Duffy and Engle-Warnick (2006) consistently show a regime change somewhere in the late 1970s or early 1980s followed by another in the mid 1980s. This result follows closely with the traditional narrative that the Federal Reserve undertook a "Monetarist experiment" during this period, wherein the Fed targeted the growth of a monetary aggregate rather than an interest rate. Using a single-equation model, Bunzel and Enders (2010) find these regimes appear in the Taylor Rule as a regime characterized by strong output gap and inflation responses (1970s) followed by a regime characterized by gradual adjustment of the federal funds rate (post 1980s). However, Estrella and Fuhrer (2003) note that these regime changes could be caused by changes elsewhere in the economy that smaller, single-equation models cannot estimate. Further, as Carvalho et al (2021) show, the estimation methodology is important to any examination of Taylor Rule coefficients. Any estimation of monetary policy is subject to an endogeneity issue, as the central bank influences and responds to

[1] Table 1 includes paper which focus specifically on estimates and breaks in the Taylor Rule. Other papers to test for structural breaks elsewhere in the model include Kim et al (2014)

[2] For examples, see the monetary policy rules in Macro Modelbase: https://www.macromodelbase.com/files/documentation_source/mmb-mprule-description.pdf?40780101f6.

changes in inflation and output. Nevertheless, Carvalho et al (2021) find that simple OLS estimates of the Taylor Rule still outperform IV estimates while still producing largely consistent model dynamics. Mavroeidis (2021) and Ikeda et al (2024) suggest addressing non-linearities in estimating VARs by using a censored and kinked VAR, and Aruoba et al (2022) suggest using a VAR with occasionally binding constraints.

Researchers also have attempted to find structural breaks in VAR models. Using a factor-augmented VAR, Bernanke and Mihov (1998) find that no simple policy variable fully captures monetary policy from 1965-1996. Instead, they find regime switches in the Fed's operating procedure in roughly 1979 and 1982, similar to the single-equation break-point models. Using structural VARs, Primiceri (2005) and Sims and Zha (2006) find similar timing of the regime changes but emphasize they are characterised by changes in the *variance* of Taylor Rule coefficients, as well as changes in the coefficient point estimates. In essence, the structural VARs suggest that monetary policy after the mid-1980s is characterized best by more consistent responses to output and inflation.

Other researchers estimated changes in the Taylor Rule using full structural models. Using an RBC model with money for 1966-1982 and 1990-2006 subsamples, Castelnuovo (2012) finds the Taylor Rule parameters are largely unchanged across subsamples, but the fed funds rate is less responsive to money growth in the second sample. Coibion and Gorodnichenko (2011) find that, while the Fed satisfied the Taylor principle in the 1970's, changes in the Taylor Rule induced determinacy during the Volcker disinflation, helping stabilize inflation. Using the same NK model and Bayesian methods as this paper, Canova (2009) finds the Fed responds more strongly to inflation after 1982, likely contributing to the Great Moderation (Stock and Watson, 2002). Using the Smets and Wouters (2007) model for 1966-1979 and 1983-2005 subsamples, Ilbas (2012) similarly finds the Fed is more responsive to inflation after 1983 as well as greater interest-rate smoothing and a lower inflation target during the

Great Moderation era. Fernández-Villaverde et al (2007) add to this body of literature by examining parameter drift, rather than clean structural breaks. They find that these deeper "structural" parameters tend to exhibit a drift over time, and parameters are likely to have substantial variation in larger samples.

3 Models

We use three benchmark macro models to estimate the Taylor Rule and test for structural breaks. For robustness, the models vary in size (small- to medium-scale) and degree of structure (few to many cross-equation restrictions).

3.1 Taylor Rules

The models contain two slightly different variants of the Taylor Rule. The VAR and NK models include a simplified version of the FRB/US Taylor Rule (equation 1),

$$i_t = \sum_{j=1}^{k} \rho_{i,j} i_{t-j} + (1 - \sum_{j=1}^{k} \rho_{i,j} i_{t-j})[\phi_\pi(\pi_t - \pi^*) + \phi_x x_t)] + \epsilon_t$$

which assumes a (suppressed) constant equilibrium nominal rate $(r + \pi^*)$ and k lags. The typical NK Taylor Rule uses $k = 1$ lag, but the VAR by construction uses $k = 2$ lags. The output gap, $x_t = (y_t - y^{POT})$, uses potential output (POT) from the Congressional Budget Office.[3] The DSGE model adds short-run feedback from the change in the output gap as Smets and Wouters (2007):

$$r_t^f = \rho_i r_{t-1}^f + (1 - \rho_i)[\phi_\pi \pi_t + \phi_y(y_t - y_t^p)] + \phi_{\Delta y}[(y_t - y_t^p) - (y_{t-1} - y_{t-1}^p)] + \epsilon_t \ , \quad (2)$$

where $r_t^f = i_t$ to momentarily sidestep notation conflict (SW use i for investment and r for the nominal rate); henceforth, i_t is the nominal rate unless noted otherwise. Equation (2) uses the DSGE concept of potential output, y^p, which denotes the level

[3] See https://www.cbo.gov/data/budget-economic-data.

that would prevail if prices were flexible and there were no markups. Estimates of the ϕ parameters provide evidence on stability of the Taylor Rule across subsamples. In contrast, Carvalho et al (2021) estimate their models with a Taylor Rule in which the Fed only targets inflation, rather than inflation and the output gap.

Neither the Taylor Rules nor the macro models incorporate UMP.[4] However, some papers have incorporated Forward Guidance (FG) into the Taylor Rule using the effective fed funds rate and FG shocks to the future policy rate as follows:

$$i_t = \rho_i i_{t-1} + (1-\rho_i)[\phi_\pi(\pi_t - \pi^*) + \phi_x(x_t)] + \epsilon_t^{MP} + \sum_{l=1}^{L} \epsilon_{l,t-l}^{I} \tag{3}$$

where $\sum_{l=1}^{L} \epsilon_{l,t-l}^{I}$ are FG shocks to the interest rate at time l, but realized at $t-l$ and ϵ_t^{MP} are the standard monetary shocks.[5] A FG shock is the difference between actual i_t and the expected rate announced by the central bank at time $t-l$. Thus, FG on future policy rates essentially extends the duration of the short-term rate at the ZLB.[6]

Although the FG-augmented Taylor Rule does not account explicitly for the quantitative easing (QE) portion of UMP, it is mathematically similar to the FG shock in the literature on QE and shadow federal funds rates. As noted by scholars from Black (1995) to Rossi (2021), the shadow rate is an option, i.e., the short-term interest rate implied by a model of the yield curve. Wu and Zhang (2019) provide a mapping of QE into a standard NK model through the shadow rate. To do this, they assume the shadow rate, s_t, follows the Central Bank (CB) balance sheet according to:

$$s_t = -\zeta(b_t^{CB} - b^{CB}) + \epsilon_t^{FG} + \epsilon_t^{MP} \tag{4}$$

[4] Because most small-scale macro models focus on the domestic economy, the Taylor Rules also do not include exchange rates, which Engel and West (2006) and Jetter et al (2019) find important. The exchange rate can, in turn, have significant effects on both inflation (through import prices and pass-through effects) and output (through net exports and competitiveness), thereby helping central banks achieve their targets for these variables. QE, in particular, involves large changes in the global supply of government bonds that may affect exchange rates and hence the policy rate.

[5] See Del Negro et al (2012), Campbell et al (2012), and Cole (2021). This specification also is called "forecast targeting" by Svensson (2017). Research with such models find a "Forward Guidance puzzle" of excessively large responses to FG news. The shadow fed funds rate controls for the effects of FG.

[6] See Section 7 for more discussion of the relationship between FG and UMP.

where ζ maps the shadow rate to the difference between bond holdings, b_t^{CB}, and their steady state level and ϵ_t^{FG} is forward guidance, and ϵ_t^{MP} is the difference between the actual and predicted shadow rates.

Figure 1 shows the shadow rate closely tracks Fed bond holdings with only three key deviations, which Wu and Zhang (2019) note coincide with the Fed's changes in FG. The early 2010 deviation coincides with the Fed signaling it would unwind its lending facilities. The 2014 decline coincides with the Fed extending its forecasted duration of the ZLB. And the early 2013 spike coincides with the so-called Taper Tantrum and is presented as a traditional monetary shock. In other words, deviations of the shadow rate from the Fed's balance sheet present themselves similarly to the FG shock.

In short, by using the shadow rate we avoid the need to include the forward guidance augmented Taylor Rule in our estimation because s_t incorporates forward guidance. Moreover, s_t also includes the effect of quantitative easing, allowing us to incorporate both aspects of unconventional monetary policy. Henceforth, we refer to the interest rate as $\hat{i}_t$ where:

$$\widehat{i}_t = \min(i_t, s_t) \tag{5}$$

to economize on notation later. The advantage of using the shadow rate is it allows for uniform comparison of the stance of monetary policy across conventional and unconventional policy periods. However, as Krippner (2020) notes, shadow rates are sensitive to their assumption of the lower bound.[7] The evolution of the shadow rate and the fed funds rate can be seen in Figure 2. The two rates are equal throughout the first subsample before diverging as interest rates hit the lower bound in 2008Q4. The shadow rate bottoms out at -2.9% in 2014. The two rates converge again once the Fed begins raising rates in 2015Q4.

[7] For robustness, we estimate each model with the DSGE-derived shadow rate of (Jones et al, 2021) and receive consistent results for each estimation. Results for this estimation are available upon request.

3.2 VAR Model

The VAR model is based on the three-variable vector, $Z_t = [y_t, \pi_t, \widehat{i}_t]'$ that includes the output gap, inflation, and the sample-specific policy rate. Abstracting from constant terms, the structural form is

$$B_0 Z_t = \sum_{i=1}^{k} B_i Z_{t-1} + u_t \ , \tag{6}$$

where the 3x1 vector of structural shocks, u_t, is identified from the Cholesky decomposition

$$B_0 = \begin{bmatrix} 1 & 0 & 0 \\ \kappa & 1 & 0 \\ (1 - \rho_{i,1} - \rho_{i,2})\phi_y & (1 - \rho_{i,1} - \rho_{i,2})\phi_\pi & 1 \end{bmatrix} \tag{7}$$

with usual diagonal covariance matrix, $\Sigma = u_t u_t'$: The ordering restrictions allow the output gap to respond only to its own innovations and hence move the slowest. Inflation responds contemporaneously to the output gap and its own innovations, while the Fed's policy rate responds to shocks in both the output gap and inflation as in the standard Taylor Rule.

This ordering identification originated with Sims (1980) but still is central to Rossi (2021) contemporary analysis. Our modest structural extension imposes interest-rate smoothing by restricting $\rho_i = \Gamma_{3,1}$, which is the (3,1) element of the first lag ($k = 1$) and second lag ($k = 2$) of the reduced-form coefficient matrices, $\Gamma_1 = B_0^{-1} B_1$ and $\Gamma_1 = B_0^{-1} B_2$.

3.3 Small-Scale New Keynesian (NK) Model

The three-equation NK model is from Clarida et al (1999) and Clarida et al (2000) and uses the same variables as the VAR. In addition to the Taylor Rule in equation

(3.1), the NK model imposes structural restrictions in the form of the IS equation and forward-looking Phillips Curve:

$$y_t = \psi[\widehat{i_t} - E_t\pi_{t+1}] + E_t y_{t+1} + \epsilon_{x,t} \tag{8}$$

$$\pi_t = \kappa y_t + \beta E_t \pi_{t+1} + \epsilon_{\pi,t} \tag{9}$$

where β is the discount factor, ψ is the coefficient of relative risk aversion, and κ is the slope of the Phillips Curve. Structural shocks $\epsilon_{y,t}$, $\epsilon_{\pi,t}$, and $\epsilon_{i,t}$ follow an AR(1) process:

$$\epsilon_{j,t} = \rho_j \epsilon_{j,t-1} + \eta_{j,t} \tag{10}$$

where $j = \{y, \pi, \widehat{i}\}$ with $0 < \rho_{j,t} < 1$ capturing the persistence of shocks and $\eta_{j,t}$ is i.i.d. with zero mean and variances σ_j^2. The model has nine parameters: six structural parameters (β, ψ, κ, ρ_i, ϕ_π, ϕ_y) and three auxiliary parameters (ρ_y, ρ_π, ρ_m).

3.4 Medium-Scale DSGE Model

The medium-scale DSGE model is from Smets and Wouters (2007), which contains the full linearized version. In addition to the Taylor Rule in equation (2), the portion of the DSGE model that most closely matches the NK model are the consumption Euler equation and expectations-augmented NK Phillips Curve:

$$c_t = c_1 c_{t-1} + (1 - c_1) E_t c_{t+1} + c_2(l_t - E_t l_{t+1}) - c_3(\widehat{i_t} - E_t \pi_{t+1} + \varepsilon_t^b) \tag{11}$$

$$\pi_t = \pi_1 \pi_{t-1} + \pi_2 E_t \pi_{t+1} - \pi_3 \mu_t^p + \varepsilon_t^p \tag{12}$$

where c_t is real consumption, l_t is hours worked, μ_t^p is the price markup, and ϵ_t^b, ϵ_t^p are structural shocks. The c_i and π_i are parameters to be estimated.[8]

[8] For a more comprehensive summary of the SW DSGE model, see Smets and Wouters (2007).

The DSGE model is more comprehensive and imposes stronger cross-equation restrictions than the NK model. For example, the NK IS Curve (8) is obtained from the simplifying assumption that $y_t = c_t$ in the forward-looking consumption Euler equation. The DSGE model does not impose this assumption but explicitly models the entire aggregate resource constraint. Similarly, the NK Phillips Curve (9) is the linearized form of the firm's simplified exogenous pricing decision. The DSGE model adds backward- looking elements into the consumption Euler equation and Phillips Curve plus a price mark-up in addition to sticky price adjustment.

The DSGE model has other advantages. It is consistent with a steady-state growth path, incorporating investment decisions and the pricing and accumulation of capital into its optimization problems. The DSGE model also has a more complex stochastic environment with seven structural shocks (productivity, technology, risk premium, spending, monetary, price-markup, and wage-markup) compared with three (demand, cost-push, and monetary), allowing richer and more flexible estimation of the effects of monetary policy. This, in turn, means the DSGE also constructs a model-consistent output gap, while the NK model requires an externally estimated output gap.

4 Econometric Specifications

Most data used in this paper come from the FRED database created by the Federal Reserve Bank of St. Louis. The VAR and NK models use: 1) the output gap constructed with the CBO's real potential GDP; 2) core PCE inflation; and 3) the short-term policy rate, $\widehat{i}_t$. The DSGE model uses the same data as Smets and Wouters (2007) but is updated and extended through 2019 and also uses $\widehat{i}_t$. The shadow federal funds rate is from Wu and Xia (2016).[9]

Following Orphanides (2001), we recognize that the distinction between real-time and revised data can be important for Taylor Rule estimation, as policymakers make

[9]See https://sites.google.com/view/jingcynthiawu/shadow-rates?authuser=0

decisions based only on data available at the time rather than on subsequent data revisions. To address this concern, we re-estimated our VAR model using real-time data and found results qualitatively consistent with our main findings using revised data. This suggests our structural break findings are robust to the choice of data vintage. Detailed real-time estimation results are provided in Appendix B.2.

4.1 Selection of Samples

Our full data sample runs from 1984:Q1 to 2019:Q4. The starting period is consistent with the literature's previous findings of a structural break during the early-1980s. We truncate the sample in 2019 to exclude the new UMP that emerged during the COVID-19 pandemic and recession.[10]

Based on the literature and conventional wisdom about known breaks in monetary policy, the subsamples are split into a Great Moderation (Subsample I) sample running from 1984:Q1 to 2007:Q2 and a Post-Crisis (Subsample II) sample running form 2007:Q3 to 2019:Q4. The beginning of the Post Crisis (2007:Q3) corresponds to the Fed's initial rate cuts and early events of the financial crisis, such as American Home Mortgage's bankruptcy, BNP Paribas noting a decline in liquidity, the Dow Jones Industrial Average's peak.

For robustness, we provide some formal evidence on the selected break points by estimating a split-sample Chow test for the VAR and testing for structural breaks (Lütkepohl, 2013). Figure 3 shows the p-value from the rolling window estimation; the horizontal dashed line indicates the 5 percent confidence level. The Chow test largely confirms our *a priori* reasoning on the subsample selection: structural breaks corresponding to the Fed's changes in operating procedure in early-1980s, as well as one near the start of the financial crisis in 2007Q3 (both indicated by the vertical dashed lines). For this reason, we continue the tradition in the literature of setting

[10]For robustness, we run a second estimation of starting in 1960 and receive results consistent with previous literature. Results for the early sample can be found in the appendix

the break periods exogenously rather than using more complicated endogenous break-point methods.

4.2 Estimation

The VAR is estimated using OLS so these Taylor Rule estimates are consistent with the recommendation of Carvalho et al (2021). The model uses $k = 2$ lags based on lag selection tests using AIC, BIC, and log likelihood criteria, which consistently selected 2 lags as optimal. Results from the lag selection test can be found in B4.[11] The structural parameters are derived from B_0 and the first own-lagged coefficient in the $\widehat{i}_t$ equation. Standard errors are obtained from the delta method (Oehlert, 1992).

Recent advances in the literature, including Mavroeidis (2021), Ikeda et al (2024), and Aruoba et al (2022) propose using VARs with altered to address nonlinearities in monetary policy estimation when rates are constrained. These methods offer a promising framework for capturing dynamics under unconventional monetary policy regimes. Thus, we first estimate our VAR using standard methods and the Wu-Xia Shadow Rate, where the nonlinearities are not present. We then estimate our model using the federal funds rate with the method in Aruoba et al (2022) to control for the lower bound in the Post-Crisis period.

The NK and DSGE models are estimated with standard Bayesian methods. Selection of priors has a significant bearing on the estimated parameters, so changing priors between subsamples can potentially bias results toward a structural break when one does not truly exist. To mitigate this bias, and for consistency with earlier research, we use the same prior distribution, mean, and standard deviation in the full sample and all subsamples: Canova's (2009) for the NK model and Smets and Wouters' (2007) for the DSGE model.[12] The likelihood function is calculated using the Kalman filter. The posterior density distribution is obtained from the calculated likelihood

[11] Quarterly estimations of VAR models often use 4 lags, and results for the model with $k = 4$ lags are in Table B6. These results are consistent with the main model specification.

[12] See Appendix A for a list of parameters, their roles in the model, and their priors.

function and prior distributions, continuing until convergence is achieved. Then the Metropolis-Hastings algorithm is used to create 2,000 draws of the posterior distribution and approximate moments of the distribution.[13] The ZLB estimation follows the method described in Cuba-Borda et al (2019) and Giovannini et al (2021) which "turns off" the monetary shock while interest rates are at the lower bound so as to avoid over-interpreting small movements in the interest rate, leading to biased results.

Following Smets and Wouters (2007), the DSGE output gap is generated from the model as the deviation from the level of output that would prevail with flexible wages and prices, y_t^p. This model-generated output gap differs from the output gap used in the VAR and NK models in two ways. First, the latter uses CBO's estimate of potential output derived from an independent growth-accounting framework.[14] Second, because CBO estimates potential output for the *full sample* it does not change across subsamples; in contrast, the DSGE output gap is estimated separately for each subsample and thus is subject to breaks in the models' structural parameters.[15]

5 Estimation Results

Tables 2 and 3 report coefficient estimates and evidence of structural breaks in model parameters for the Taylor Rule and other non-policy coefficient estimates, respectively, in the full sample (Full) and each subsample (I-II). There are two subsample II periods depending on which funds rate is used: II$\widehat{i}$ (shadow funds rate) and IIi (fed funds rate). The tables also include coefficient changes between subsamples and, for the VAR model, significance of the t-tests for differences. Coefficient magnitudes may vary across models due to differences in model variables and structure, and thus should be compared mainly across subsamples within models.

[13] Estimation is performed using a modified version of Johannes Pfeifer's dynare code for Smets and Wouter's model. Pfeifer's code can be found at https://github.com/JohannesPfeifer/DSGE_mod/tree/master/Smets_Wouters_2007, and Dynare can be downloaded at https://dynare.org.

[14] See Shackleton et al (2018).

[15] Quast and Wolters (2023) discusses the reliability and inconsistencies of output gap estimation across samples as data changes.

5.1 Taylor Rule Parameters

During the Great Moderation, the estimated coefficients (Table 2, column I) are broadly consistent with the prior literature.[16] The Fed tends to respond strongly to inflation when setting interest rates ($\phi_\pi \geq 1.97$ in all models during this period). Interest rate persistence is similar across models but a bit lower in the NK model (approximately .6 versus .8). The DSGE's response to output growth ($\phi_{\Delta y}$) is significant.

After the GFC, the estimated coefficients (column IIi) tended to see an economically meaningful change. In all three models, the Fed became less responsive to inflation as ϕ_π declined by economically large amounts. Changes in other coefficients generally were smaller in absolute value and less systematic and significant. The response to output (ϕ_y) declined significantly in the VAR (-1.66) but increased insignificantly in the NK model (.25) and in the DSGE model (.04). While these changes aren't consistently significant, this divergence between model estimates reflects differences in identification assumptions, with the VAR using ordering restrictions to identify backward-looking relationships while the NK and DSGE models impose cross-equation constraints from forward-looking optimizing behavior. Persistence (ρ_i) was essentially unchanged in all 3 models ($-.01$ in the VAR, $-.01$ in the NK model, and $-.01$ in the DSGE). The Fed's response to output *growth* ($\phi_{\Delta y}$) was unchanged.

Perhaps surprisingly, coefficient estimates with the actual fed funds rate (IIi) are quite similar to those using the shadow rate (II$\widehat{i}$), as shown in Table 2. In fact, the coefficient magnitudes are essentially the same statistically with only a few key exceptions. In the VAR model, ϕ_π is positive but still not significantly different from the II$\widehat{i}$ estimates. In the NK model, ϕ_y is smaller in the shadow rate estimation (.64 versus .94).

[16] Specifically, the Great Moderation point estimates for ϕ_π are consistent with those estimated in Carvalho et al (2021) via both OLS (2.75) and IV (2.63).

Changes in estimated Taylor Rules during period II show a decline in ϕ_π relative to ϕ_y for both the shadow and fed funds rates in the NK and DSGE models. This result could reflect changes in the preferences of monetary policy makers, as was suggested for previous breaks in Taylor Rule coefficients.[17] Bordo and Istrefi (2023) find the FOMC shifted toward being dominated by Doves (higher weight on the output gap) after the GFC (period II). Kocherlakota (2019) goes further, arguing that policy makers have private information about their objectives (which may be influenced by non-economic factors) that only affects economic outcomes through the policy choice and thus acts like a taste shifter. If so, the unconditional independence of policy rules assumed in the benchmark macro models would be violated. Time-varying preferences could be modeled with ϕ_π and ϕ_y as functions of FOMC composition over time.

5.2 Non-policy Parameters

Table 3 reports estimates of the models' non-policy parameters. Note that, unlike the Taylor Rule parameters, the NK and DSGE coefficients are not directly comparable due to substantial differences in the size and structural restrictions of the two models. Nevertheless, these coefficient estimates during the Great Moderation (column I) are generally consistent with the literature. In the VAR model, the slope of the Phillips Curve is not statistically significant. In NK model, the slope of the IS Curve (ψ) is negative and small in absolute value, implying a relatively high coefficient of relative risk aversion of about 30. The slope of the Phillips Curve (κ) and the expectations feedback are both positive and relatively high but significantly less than 1.0. In the DSGE model, there are too many parameters to discuss individually. However, these DSGE coefficient estimates, along with those in period I, are roughly in line with what

[17] Breaks leading into the Great Moderation (period I) were described as a shift in the preferences of FOMC members toward favoring inflation stability by Canova (2009), Castelnuovo (2012), Ilbas (2012), and Lakdawala (2016), for examples. However, Debortoli and Nunes (2014) and Scott (2016) caution against interpreting structural shifts in the policy rule as simply a change in preferences, noting that shifts in the policy rule can obscure differences between factors inside and outside a policy maker's control.

has been reported in Smets and Wouters (2007) and subsequent estimates of their model.

After the GFC, many of the estimated non-policy coefficients (column II$\widehat{i}$) in the NK and DSGE models exhibit significant changes. In the NK model, the IS Curve slope (ψ) became more negative ($-.2$ versus $-.03$). In turn, the coefficient of relative risk aversion fell to about 5, and the output gap is more sensitive to the real rate. The Phillips Curve slope (κ) declined considerably (.38 to .05). Inflation expectations feedback (β) also decreased significantly (.91 to .70).

Similar to the Taylor Rule coefficients, estimated non-policy coefficients are quite similar whether the model was estimated using the shadow rate (II$\widehat{i}$) or the actual fed funds rate (IIi). Specifically, the slope of the IS and Phillips Curves is nearly identical across estimations (-.19 vs -.2 and .03 vs .05, respectively). Only inflation expectations feedback meaningfully changes, as inflation expectations have a larger effect in the shadow rate estimation (.82 vs .70).

In the DSGE model, several coefficients changed notably after the GFC. Two long-run coefficients, steady state growth ($\overline{\gamma}$) and hours ($\bar{l}$), fell by economically and statistically significant amounts (.48 to .22 and .79 to -4, respectively). In contrast, steady state inflation ($\overline{\pi}$) essentially was unchanged. The capital share (α) declined by almost half and the external habit (λ) increased notably (.52 to .63). The DSGE model also provides an estimates the natural real interest rate as a function of the rate of time preference, β, and risk aversion, σ_c.[18] Estimates for period I are larger than many in the literature but similar to the original estimates in Smets and Wouters (2007). In period II, the real rate estimates fell almost in half (3.1 to 1.7 percent). The remaining DSGE coefficients did not change statistically significantly.

Notably in the DSGE estimation, *none* of the non-policy parameters exhibit statistical or economic differences between the shadow rate or fed funds rate estimations

[18]The natural real interest rate is calculated as in Smets and Wouters (2007): $\bar{r} = (\frac{\gamma^{\sigma_c}\Pi}{\beta} - 1)100$.

(II$\widehat{i}$ vs IIi). In turn, the differences noted between period I and and period IIi are still present: steady state growth ($\bar{\gamma}$) and hours ($\bar{l}$) decline while steady state inflation ($\bar{\pi}$) remains constant. Additionally, external habit λ increases and capital share α increases. As a result, the implied real interest rate is still lower than the subsample I value.

Overall, the estimated changes in the DSGE coefficients, especially during period II, reinforce the need for macro models to incorporate time variation to properly identify monetary policies. Steady state (trend) growth varied, perhaps due to breaks in productivity trends (e.g., Fernald (2014)) and variation in the marginal product of capital. Our estimates of the (typically fixed) natural real interest rate (r) varied too, similar to Del Negro et al (2019).[19] The benchmark models also assume a fixed inflation target, $\bar{\pi}$, but inflation volatility during the Great Inflation and the subsequent "opportunistic disinflation" (Orphanides and Wilcox, 2002) suggest the target also may be time varying. Our estimates of the slope of the NK Phillips Curve and degree of nominal price and wage stickiness changed, as in Kim et al (2014) and Jorgensen and Lansing (2021), for examples, suggesting a need to model variation in the underlying related trends. Estimated increases in coefficients on inflation expectations may reflect increasing efficiency of monetary policy, similar to the finding in Anzuini and Rossi (2022). Finally, consumer risk aversion also declined, but it is hard to draw hard conclusions about the cause(s) given the challenges in estimating this parameter (Calvet et al, 2021).

[19] As noted earlier, our estimates of the natural real interest rate are consistent with Smets and Wouters (2007) original estimate of r^* but are considerably higher than traditional estimates in the literature. Holston et al (2017) estimated the natural rate of interest to be close to zero after the financial crisis, Del Negro et al (2019) estimate it to be slightly above one. Further, Levrero (2021) notes that estimates of the natural rate can vary widely based on the definition of the natural rate and the estimation method used. Resolving these large inconsistencies around r^* is key to explaining structural breaks. Time-variation in the natural rate might follow Laubach and Williams (2003) and Bjørnland et al (2011), whose natural real rate of interest, $r^* = (1/\sigma)\bar{\gamma} + \beta$, and its law of motion, $r_t^* = c\bar{\gamma}_t + z_t$, would be added to the Taylor Rule. Steady state growth varies over time ($\bar{\gamma}_t$) while z_t captures other determinants of r_t^*, such as household rate of time preference. See also Hamilton et al (2016) for long-run variation in r_t, which may differ from r^* but likely has similar trend variation.

5.3 Discussion

This section presents evidence of structural breaks in many macro model coefficients. However, the analysis cannot verify whether UMP is responsible for the observed breaks without a clear alternative model that includes equations for UMP, as discussed in Section 7. Nevertheless, it is instructive to summarize the results thus far and assess their implications for monetary policy.

Three broad conclusions can be deduced from the results:

- *Parameter types* – Structural breaks occur in both Taylor Rule (policy) and non-policy coefficients, making it more difficult to isolate the effects of omitted policies on the Taylor Rule. Because the policy and non-policy parameters are estimated jointly, changes in the latter can influence estimates of the former.
- *Estimation Method* – Structural breaks are consistent between models estimated with the ZLB-specific fed funds estimation and with the shadow rate.
- *Models* – Structural breaks are not exactly comparable between the small models (VAR, NK) and the larger DSGE model. While none of them includes UMP, the DSGE model has more variables that are likely influenced (directly or indirectly) by UMP. It is difficult to identify whether coefficient changes, especially non-policy, are due to omitted monetary policies or to changes in the private-sector economic structure.

Thus, the evidence presented thus far does not conclusively indicate whether UMP is responsible for the comprehensive and heterogeneous structural breaks in model parameters. However, the collective effects of structural breaks are reflected in the policy rate estimates for each sub-sample shown in Figure 4. Using the NK model, the figure plots the in-sample fitted values and out-of-sample values of the shadow rate (equal to fed funds when it's above the ZLB) against the actual data for each sub-sample (thick black line). It uses actual post-sample data (2020-2022) to project the shadow rate during the COVID-19 Pandemic to help assess the likelihood of additional changes to monetary policy.[20]

Three key results emerge from Figure 4. First and not surprisingly, the models fit the shadow rate best during the sub-sample in which their parameters were estimated. Post-crisis (Wu-Xia) estimates stem from data with UMP (though not equations in the model), whereas Post-Crisis (FF) estimates were not influenced by UMP yet yield similar estimates. In fact, the Post-Crisis (FF) regime fits the data slightly better during period II. A third and notable result is the gaps between all projected shadow rates and actual data after 2020 suggest that recent monetary policy was more stimulative than previous policies would have been. The Post-Crisis (FF) regime would have cut and raised the shadow rate sooner faster than observed, and the rate would have been about 1 percentage point higher by the end of 2022. The Great Moderation and Post-Crisis (Wu-Xia) regime, which featured consistently low and stable inflation, would have been significantly tighter: the shadow rate would not have breached the ZLB and it would have been at least double the realized rate at the end of 2022. Together, these three results suggest three hypotheses: 1) monetary policy may have shifted again in 2020, perhaps with the introduction of average inflation targeting; 2) UMP may have been more influential during the Pandemic; and 3) post-Pandemic monetary policy may bear considerable responsibility for the rise in inflation to about four times the Fed's target (2 percent).

6 Additional Diagnostics

The previous section documents evidence of significant breaks in the individual parameters of the models. Motivated by the evidence, this section characterizes the collective effect of breaks in model parameters using two additional diagnostic measures: 1) estimated structural shocks and 2) dynamic responses to structural shocks[21]. Changes in

[20] Attempts to estimate the NK model through 2022 (instead of 2019) were unsuccessful. This challenge suggests policy and/or non-policy structure may have changed considerably during this time.

[21] A third measure is estimated output gaps, which are discussed furhter in the appendix.

these measures reveal the macroeconomic implications of parameter breaks in period II, and providing supporting evidence of structural breaks and a fuller understanding of the nature of the observed changes.

6.1 Structural shocks

The time series characteristics of each model's estimated structural shocks provide one way to summarize the comprehensive impact of parameter breaks. Figure 5 plots the estimated monetary policy shocks for each model.[22] Notably, for the IIi estimation, the monetary shock is "turned off" from 2008Q4 until 2015Q4 in the NK and DSGE, as the fed funds rate was against the ZLB (Cuba-Borda et al, 2019).

In the VAR, the variance of the monetary shock stayed roughly the same between subsamples I and II$\widehat{i}$. In contrast, the variance of the shock decreases substantially in the NK model and increased in the DSGE model. While it is difficult to compare coefficient magnitudes across models, this difference across models' shock structure is worth noting. Put simply, the nature of the shocks is changing across models.

The shock properties of the IIi estimation are limited to periods when the fed funds rate is away from the ZLB in the NK and DSGE, while the II$\widehat{i}$ estimation is not bound by the ZLB and accounts for the full subsample. As such, comparing the IIi and II$\widehat{i}$ shocks are instructive as the II$\widehat{i}$ estimation reveals shocks from the Fed's non-interest rate policies while the IIi estimation holds these shocks to zero. Across both the NK and DSGE models, the variance of the monetary shock is smaller in the IIi estimation.

Changes in the time series properties of the estimated monetary shocks indirectly reflect the effects of changes in the estimated coefficients of Taylor Rule and non-policy structural equations reported in Section 5. While the estimated model coefficients

[22] See Figures C2 and C3 in the Appendix for plots of other structural shocks in the NK and DSGE models.

exhibit various breaks, the overall time series properties of the monetary shocks in period II reveal moderate changes in variability and persistence across subsamples.[23]

6.2 Impulse Responses

Changes in the Taylor Rule and non-policy parameters also affect the dynamic properties of the macro models. Figure 6 shows impulse responses to a 100-basis-point shock to the federal funds rate. The responses are broadly consistent with prior evidence for each model and, with few exceptions, *qualitatively* similar across models and samples. Monetary tightening produces a familiar, modest decline in output and inflation, followed by a slow return to steady state for about 1-3 years. The funds rate paths are nearly the same, decaying slowly from 100 basis points in a similar fashion across models with only modest differences in the degree of persistence. This result is consistent with the finding in Carvalho et al (2021) that different estimation methods provide largely unbiased impulse response functions, although estimation methods may vary in precision.

However, the output gap and inflation responses exhibit somewhat larger and more economically important *quantitative* differences across models and subsamples. For example, although the average output response is similar across models and samples, the absolute magnitude of output responses varies more across samples in the VAR and NK models than the DSGE model. Also, the DSGE has notably larger (in absolute value) and economically different inflation responses than the other models. A lack of consistency across models perhaps is to be expected given their different sizes and restrictions, but the relative inconsistency of subsample responses across models is striking. That is, the largest absolute response for each model is not associated with the same subsample.

[23] The time series properties of the other estimated NK and DSGE structural shocks also exhibit a variety of changes but there are too many to discuss here. See the Appendix for more details and discussion.

Although subsample heterogeneity across responses may be providing additional evidence of structural breaks, most differences are economically moderate for at least two reasons. First, as noted in Section 6.1, the variances and persistence of the structural shocks change considerably across subsamples, which also impact the coefficient estimates. Unlike the monetary shock fixed at 100 basis points, impulse responses based on shocks' estimated standard deviations (not displayed but available upon request) vary much more. Second, data-consistent dynamics are the inherent goal of model estimation. Thus, while breaks in the economic structure may occur in some coefficients (e.g., the Taylor Rule), offsetting breaks in other parameters (e.g., non-policy) may occur simultaneously to maintain dynamic properties consistent with the data.

To better understand the effects of structural breaks in Taylor Rule parameters, we conducted a counterfactual exercise in which the non-policy parameters are held fixed at their full-sample estimates. Figure 7 shows impulse responses to a 100-basis point fed funds shock using models in which only Taylor Rule coefficients change across subsamples, thus better illustrating the effects of structural breaks in policy on model dynamics.

The counterfactual responses reveal three important insights. First, absolute magnitudes responses are roughly one-third the size of the unrestricted responses (Figure 6) for all but the DSGE output gap response, which nearly doubles. Second, the counterfactual responses are much more consistent across subsamples with smaller qualitative differences. Third, the Great Moderation (period I) responses more consistently differ from both post-Crisis (period II) responses, which are similar to each other. The DSGE fed funds rate responses exhibit a short-lived amplification after the shock while the NK dynamics are similar to the unrestricted responses. Inflation is more responsive in the Great Moderation in the VAR. Overall, these counterfactual responses show that breaks in the non-policy parameters mute the volatility of responses differing only in Taylor Rules. Changes in Taylor Rule coefficients across samples and models thus have limited effects on model dynamics.[24]

To summarize the dynamics results, changes in model coefficients have modest economic effects on dynamics when all coefficients are allowed change. Changes in subsets of the coefficients alter dynamic responses by magnitudes that are larger and economically more important, but these effects largely offset when all model coefficients are allowed to change.

7 Explaining Structural Breaks

Economically and statistically significant structural breaks in both Taylor Rule *and* non-policy coefficients makes inference about cause(s) of the breaks much more difficult. If breaks had occurred only in the Taylor Rule, it might be possible to discern shifts due to UMP using the shadow rate as an approximate control. However, with many key non-policy coefficients exhibiting economically significant time variation in point estimates and variances, it does not appear feasible to identify breaks induced by UMP separately from breaks unrelated to policy in the benchmark models. Thus, the benchmark macro models may be exhibiting structural breaks in parameters that actually reflect some or all of these sources of time variation rather than breaks associated with UMP.

If so, then a substantially more complex macro model(s) is needed to account for the observed time variation. The work of Fernández-Villaverde et al (2007) and Fernández-Villaverde and Rubio-Ramírez (2010) is an important step in this direction and good launching point for introducing time variation. However, even that work does not include all sources of observed non-policy structural breaks discussed earlier, and it does not include any elements of UMP. Incorporating all or even some of these

[24] See the appendix for impulse response functions using the full sample Taylor Rule and subsample non-policy coefficients.

extensions and estimating the expanded models is a challenging task that is beyond the scope of the current paper.

Ultimately, future research will need to address the omission of UMP by explicitly specifying the actual policy rule(s) governing UMP and the non-policy structure through which it transmits to macroeconomy, such as long-term interest rates. Below, we give a high-level summary discussion of the issues pertaining to each of these additions.

Forward Guidance (FG) – Developed during the (relatively) low-interest rate period of the early 2000s, FG was tested first during the subsequent increase of the federal funds rate in 2004 (Gürkaynak et al, 2004). The main implementation of FG occurred during the GFC when the federal funds rate hit the ZLB for six years. Rather than using the shadow funds rate, it may be necessary to insert the prototypical FG model (equation 3) into the benchmark macro models.

Quantitative Easing (QE) – From 2008-2014, the Fed substantially expanded its Open Market Operations (OMO) to conduct large-scale asset purchases (LSAP) of: 1) mortgage-backed securities, to ease bank risk and lower mortgage rates; and 2) longer term Treasury bonds, to increase maturity and lower long-term risk-free rates. This QE strategy added a new monetary policy governing the conduct and objective(s) of LSAPs (and eventually sales, or LSAS), as well as a more explicit specification of the banking sector and its interaction with the economy. One manifestation of QE policy is a simple balance sheet rule(s) that emulates the Taylor formula, such as those proposed in Sims et al (2021) and Sims and Wu (2021):

$$B_t = \rho_B B_{t-1} + (1 - \rho_B)[\theta_\pi(\pi_t - \pi^*) + \theta_x x_t] + \nu_t \tag{13}$$

where B_t is the Fed's holding of long-term bonds.[25]

[25] In practice, the FOMC appears to implement such a rule as *changes* in Fed's target purchases of QE securities. See the regular FOMC statements during and after the two recent periods of ZLB (2008-2015 and 2020-2022) for details.

The other manifestation is a target(s) and rule(s) for Fed mortgage and/or Treasury real bond holdings and related long-term interest rates. While the Fed does not explicitly announce a target long-term rate, the balance sheet rule implicitly suggests one. In Vázquez (2009) and Carlstrom et al (2017), the Taylor Rule for the short rate becomes:

$$i_t = \rho_i i_{t-1} + (1-\rho_i)[\phi_\pi(\pi_t - \pi^*) + \phi_x x_t) + \phi_{tp} tp_t] + \epsilon_t^{MP} \tag{14}$$

where tp_t is the term premium. For robustness, we estimate the VAR using the Taylor Rule specified in (14), and full results can be found in Appendix B and Table B3. Overall, we find consistent results with our standard estimation, even with this inclusion of the term premium. In turn, macro models most likely need to introduce explicit specifications of QE policies and asset-pricing equation(s) to identify the effects of UMP properly.[26]

Expanded Liquidity Facilities (ELF) – During and after the Financial Crisis, the Fed developed new policy tools to provide liquidity and improve functioning of financial markets. These policies, such as the interest rate on reserve balances (IORB) and the Term Asset-Backed Securities Loan Facility (TALF), marked a drastic change in the Fed's policy implementation and crisis management, respectively. However, because most macro models are focused on a lower frequency, these policies are unlikely to alter the Fed's objectives and, in turn, the Taylor Rule.

To recap this section, observed structural breaks in the Taylor Rule may reflect the effects of omitting variables and equations associated with UMP. If so, expanding the macro models to incorporate the UMP and related non-policy equations may be necessary to fully and properly capture the effects of UMP. Recent research is

[26]Modeling UMP became even more challenging in 2020 with two new responses to the COVID-19 pandemic: 1) expanded QE that included purchases of investment-grade corporate bonds via the Secondary Market Corporate Credit Facility (SMCCF) and short-term state and local government notes via the Municipal Liquidity Facility (MLF); and 2) new direct lending to small and medium-sized businesses via the Paycheck Protection Program Liquidity Facility (PPPLF) and the Main Street Lending Program.

developing theoretical foundations for some types of UMP.[27] However, no theoretical model includes all elements of UMP, and there is little or no estimation of such models. Addressing these deficiencies is important for future research.

8 Conclusions

Three classes of benchmark macroeconomic models exhibit economically and statistically significant breaks in their Taylor Rule and non-policy coefficients after 2007:Q3. Evidence of breaks is stronger and more widespread in the larger DSGE model. The main result pertaining to the Taylor Rule is a decline in the strength of the Fed's response to inflation relative to its response to output, making the Taylor Rule somewhat more similar to its form in the period before the Great Moderation. A structural break(s) was likely given the implementation of UMP that are not included explicitly in the benchmark models. However, it is unclear whether these widespread and heterogeneous breaks reflect the effects of UMP or something else. And, perhaps surprisingly, using a shadow rate to control for UMP and avoid the ZLB does not alter the estimation outcomes much.

The observed structural breaks are heterogeneous and challenging to interpret well. One complicating factor is that breaks in non-policy coefficients influence the models as much or more than breaks in the Taylor Rule coefficients. Thus, many elements of the benchmark models may be susceptible to time variation that is not included in them. The first important task is to build and estimate a macro model(s) that incorporate some or all of the time-varying elements that are clouding inference about the effects of UMP. Then testing the revised model for structural breaks is more likely to identify the effects of UMP.

A second complicating factor is that the benchmark macro models do not include explicit specifications of UMP. Consequently, the observed structural breaks may be

[27] Examples include Gertler and Karadi (2011), Bauer and Rudebusch (2014), Hagedorn et al (2019), and Sims and Wu (2021).

simply reflecting the estimation effects of omitted variables (and equations) rather than UMP. A form of Lucas (1976) also may be at work. After controlling for potential time-variation in macro models, the obvious remedy is to include explicit specifications of FG (augmented Taylor Rule or more), QE, and possibly ELF into the model(s). Testing for structural breaks in the revised model's non-policy block of equations should more effectively identify the effects of the introduction of UMP.

Neither the task of controlling for time-variation in macro models nor the task of introducing explicit UMP instruments and rules is easy or fast. However, both are potentially important directions for future research and analysis of modern monetary policy.

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Declarations

No funding was received for conducting this study.

Tables and Figures

Fig. 1 The Wu-Xia Shadow Rate and the Fed's bond holdings

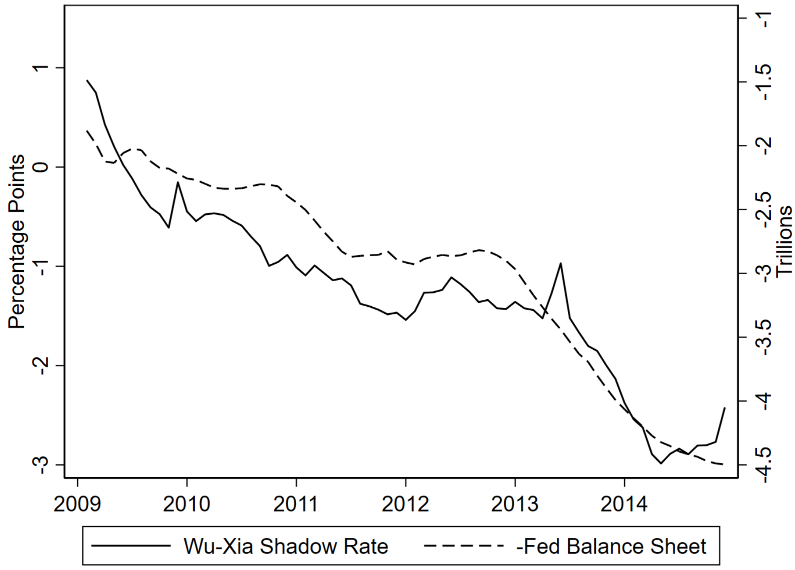


Fig. 2 The Wu-Xia Shadow Rate and the federal funds rate

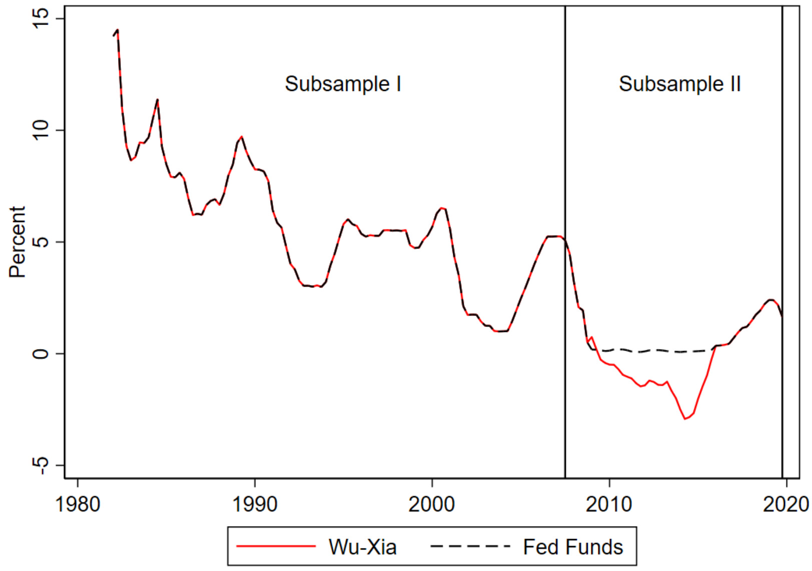


Fig. 3 P-value from endogenous breakpoint Chow test

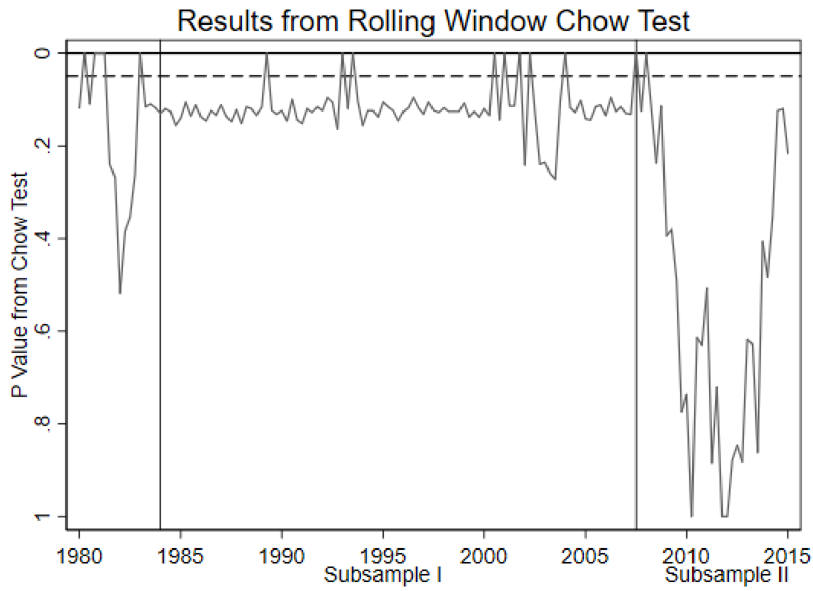


Fig. 4 NK Model Fitted and Projected Shadow Rates

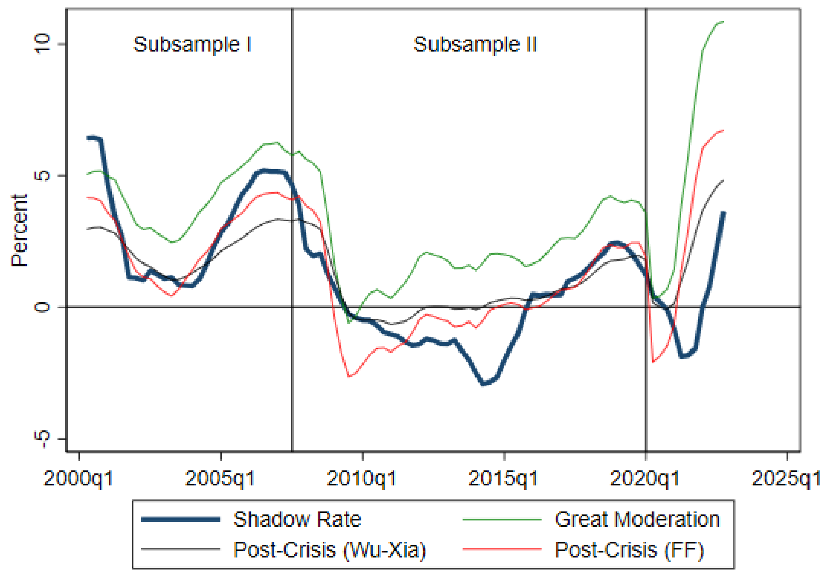


Fig. 5 Structural monetary shocks by model and sample

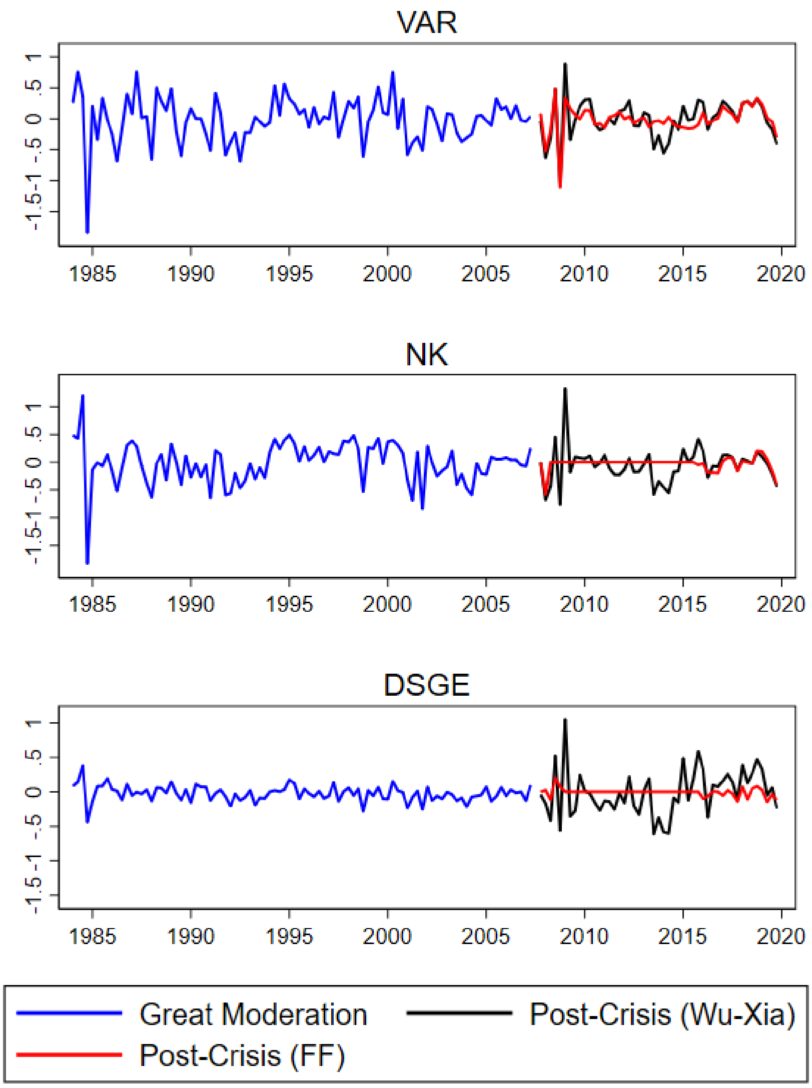


Fig. 6 Impulse response to a 100bp monetary shock by model and sample

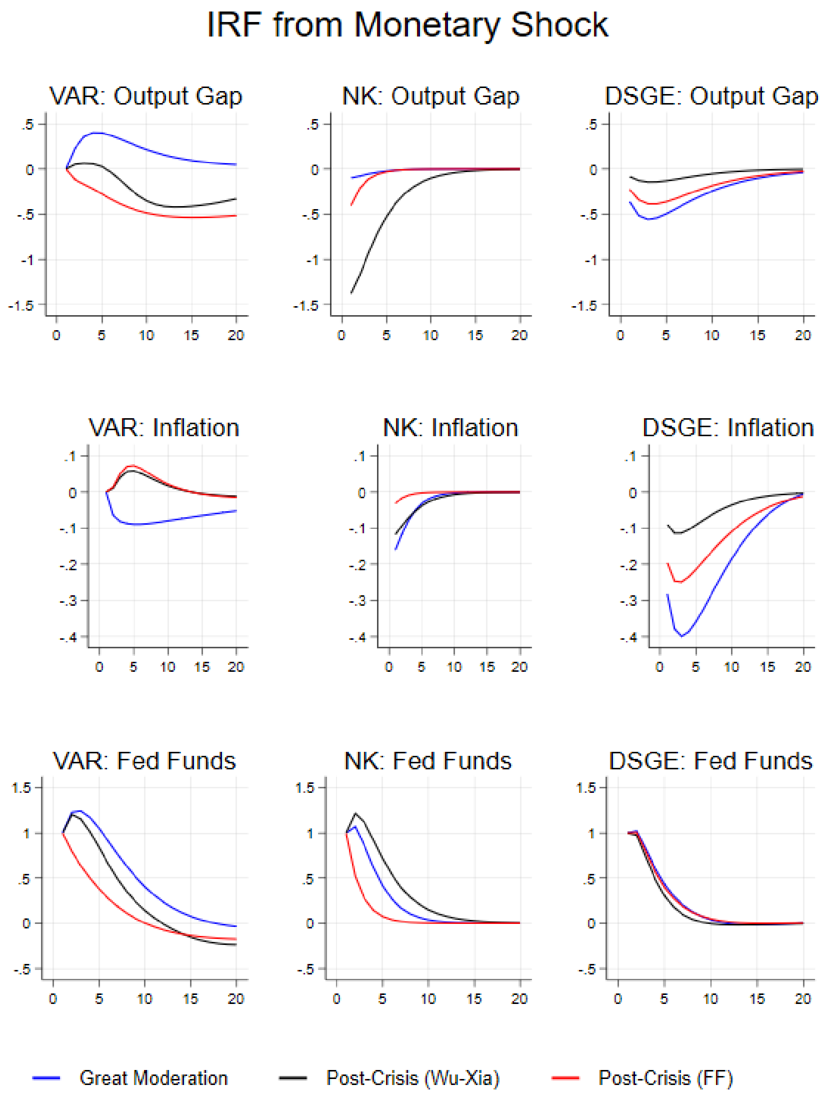
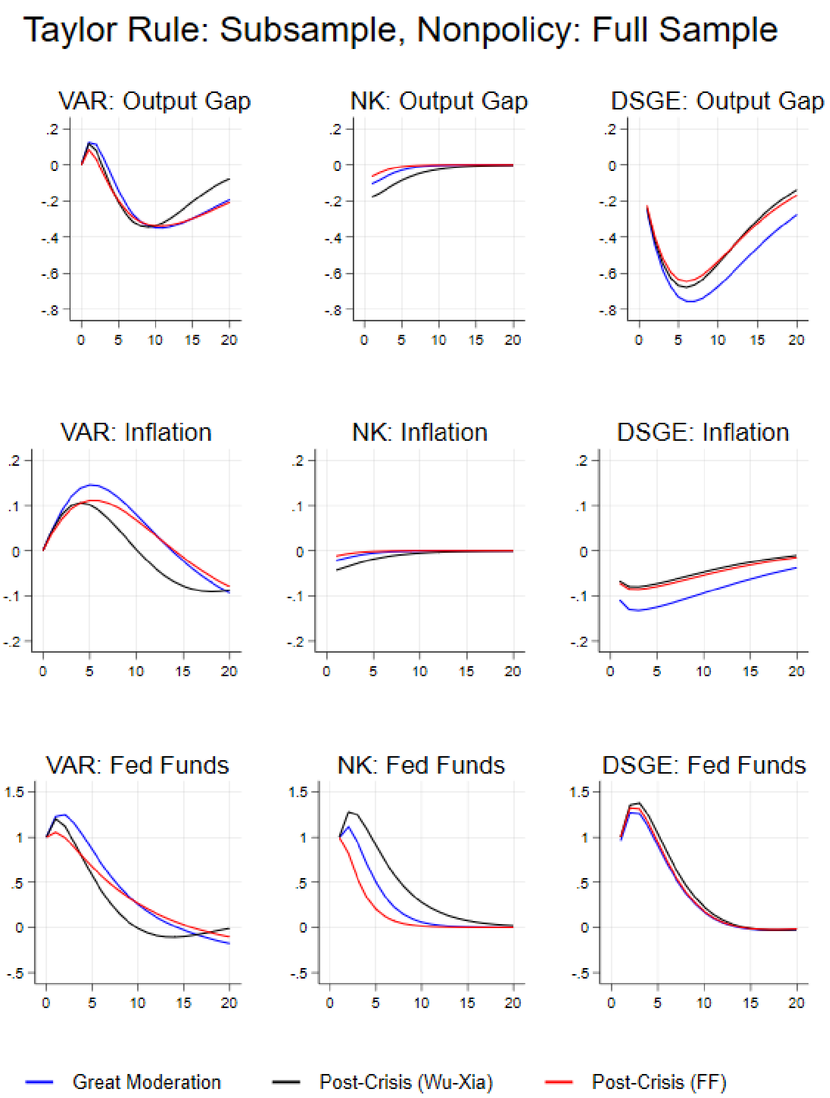


Fig. 7 Counterfactual responses to a 100bp monetary shock by model and sample



Appendix A Bayesian Estimation Priors

The prior distributions, means, and standard deviations used in estimation, given in Table A1, are similar to those used in Canova (2009) for the NK model. The slope of the

IS curve, ψ, and Phillips Curve, κ, have gamma distributions with a prior mean of -.5 and 1, respectively. The inflation feedback parameter, β, has a beta distribution and a prior mean near a rational expectations benchmark at .98. The monetary parameters are set at $\rho_i = .8, \phi_y = .5$ and $\phi_\pi = 1.3$. Additionally, the prior for the inflation parameter, ϕ_π is truncated at 1 to not allow indeterminacy.

For the DSGE model, we use the same priors as Smets and Wouters (2007), given in Table A2. The time preference rate is set at 0.25 (corresponding to $\beta = .9975$), the steady state inflation ($\bar{\pi}$) and growth rate ($\bar{\gamma}$) are set at 0.62 and and .4, respectively (corresponding to an annualized 2.5% inflation rate and 1.6% real growth rate). Steady state hours, $\bar{l}$, is set at 0, Firsh elasticity, σ_l, is 2, while risk aversion, σ_c, and habit formation, λ, are set at 1.5 and .7, respectively. The Calvo parameters, ξ_p and ξ_w, are both .5, and wage and price indexation, (ι_p, ι_w) are also .5. Finally, capacity utilization, ψ, is set at .5, and the fixed cost and capital shares (Φ and α) are 1.25 and .3, respectively. The monetary autoregressive parameter, ρ_i, is set at .75, and the monetary feedback parameters, $\phi_\pi, \phi_y, \phi_{\Delta y}$ are set at 1.5, .12, and .12, respectively, Similar to the NK model, the prior for ϕ_π is truncated at 1 to require determinacy.

Appendix B Robustness Check Estimations

B.1 Carlstrom et al (2017) Taylor Rule Estimation

We estimate an augmented Taylor Rule, similar to the one in Carlstrom et al (2017) that includes the term premium to better capture unconventional monetary policy effects. This specification modifies our baseline VAR Taylor Rule to:

$$i_t = \sum_{j=1}^{2} \rho_{i,j} i_{t-j} + (1 - \sum_{j=1}^{2} \rho_{i,j})[\phi_\pi(\pi_t - \pi^*) + \phi_x x_t) + \phi_{tp} tp_t] + \epsilon_t^{MP} \tag{B1}$$

where tp_t is the term premium spread between the 10-year Treasury yield and the federal funds rate. Table B3 reports coefficient estimates and evidence of structural breaks in the augmented Taylor Rule for the full sample and each subsample.

Comparing the Carlstrom et al (2017) specification to our baseline estimation reveals that the coefficients are fairly similar between the two estimations. During the Great Moderation (column I), the inflation response coefficient (ϕ_π) is largely similar at 3.27 compared to 2.87 in the baseline estimation. The output response coefficient (ϕ_y), however, is notably lower at 1.51 versus 2.93 in the baseline. The autoregressive properties ($\rho_{i,1}$ and $\rho_{i,2}$) are quite similar between specifications, with the autoregressive coefficients at .88 vs .89 in the baseline.

A key finding is that the policy rate responds strongly to shifts in the term premium across all samples and subsamples. The term premium coefficient (ϕ_{tp}) is positive and economically large in the full sample at 2.32, and even higher during the Great Moderation at 2.52. This indicates that throughout both conventional and unconventional policy periods, the Fed's policy rate has responded systematically to movements in the term premium. After the GFC, important differences emerge compared to our baseline results. The decline in the inflation response coefficient is greater in the Carlstrom estimation than in the main estimation. Using the shadow rate (II$\hat{\imath}$), ϕ_π falls to 1.27, representing a decline of -2.0 compared to -1.39 in the baseline estimation. This larger magnitude suggests that when term premium dynamics are explicitly controlled for, the apparent reduction in the Fed's inflation responsiveness becomes even more pronounced.

The output coefficient exhibits markedly different behavior in the Carlstrom specification. Using the shadow rate estimation, the output response shows essentially no change between the Great Moderation and post-crisis periods, with ϕ_y moving from 1.81 to 1.71 — not a statistically significant change. This is a notable departure from the main estimation, which showed a significant decline of -1.53 in the output response.

However, the decline in the output coefficient is present in the fed funds estimation (IIi), where ϕ_y falls by -0.75, more consistent with the baseline results. The autoregressive properties remain similar across all estimations, showing little meaningful change between subsamples.

The sensitivity of the policy rate to the term premium varies quite substantially between subsamples. In the Great Moderation to shadow rate comparison, the term premium coefficient declines by -0.93, from 2.52 to 1.59, which is statistically significant. Meanwhile, the fed funds estimation shows a smaller decline of -0.57, from 2.52 to 1.98, which is not statistically significant. This difference is particularly noteworthy because the Wu-Xia shadow rate is calculated using longer-term interest rates, including the 10-year Treasury rate that comprises our term premium variable. This mechanical relationship may explain why the shadow rate estimation shows a larger decline in term premium sensitivity.

B.2 Real-Time Data

Table B5 reports Taylor Rule parameter estimates using real-time data. Following Orphanides (2001), who demonstrated the importance of using real-time data in estimating monetary policy rules, these results serve as a robustness check on our main findings in Table 2.

During the Great Moderation (subsample I), the real-time estimates show lower coefficients compared to Table 2. The Fed's response to inflation (ϕ_π) is 1.51, substantially lower than the 2.67 estimated with revised data. Similarly, the output response (ϕ_y) is 2.20, compared to 3.10 in the baseline specification. Interest rate persistence (ρ_i) is similar across specifications at approximately 0.9.

After the GFC (subsample II), the real-time estimates continue to show lower coefficient levels. The inflation response declines to negative values in both specifications, while the output response remains negative but smaller in magnitude. The persistence coefficient remains relatively stable.

The structural break patterns, however, are consistent with our main findings. Both the inflation response coefficient (ϕ_π) and output response coefficient (ϕ_y) decline significantly between subsample I and subsample II. The decline in ϕ_π is particularly pronounced, falling by .95 and 1.14 in the two specifications—magnitudes comparable to our baseline results in Table 2.

The differences in coefficient levels likely reflect the greater quarter-to-quarter variability inherent in real-time data. This higher volatility is evident in the autoregressive properties of the data: the lagged output gap coefficient is 0.83 in the real-time VAR compared to 0.93 in the baseline specification, while the lagged inflation coefficient declines dramatically to 0.2 from 0.93. This reduced persistence makes real-time parameter estimates more difficult to estimate with precision, as reflected in the wider confidence intervals. Despite these estimation challenges, our core finding of structural breaks in Taylor Rule parameters after 2007 remains robust to the use of real-time data.

B.3 Additional Lags Estimation

When selecting our lags for the VAR model, we use statistical information criterion to determine the optimal number of lags. AIC, BIC, and HQIC each point to $k = 2$ lags as optimal, as shown in Table B4. However, it is common with quarterly data to use 4 lags to capture a year's worth of lags. As such, we also run our VAR model using $k = 4$ lags. Results for this specification are in table B6.

Point estimates of the coefficients tend to be somewhat smaller with $k = 4$ lags, though the change in the coefficients remains similar. The coefficient on inflation, ϕ_π is 1.99 for the full sample, compared to 1.74 in the main estimation. The sum of the autoregressive coefficients is somewhat larger with 4 lags, at $\sum_{j=1}^{4} \rho_{i,j}$ =.94, vs .89

with two lags. The output gap response, ϕ_y is is somewhat smaller at 2.91 for the full sample, compared to 3.40 in the main estimation. During the Great Moderation, the Fed's response to both inflation and the output gap are somewhat lower than in the main estimation, at 2.54 and 2.91, respectively, vs 2.87 and 3.25. The sum of the autoregressive coefficients is somewhat lower as well, at .84 with 4 lags against .89 with two lags.

After the GFC, both ϕ_π and ϕ_y decline in both the II$\widehat{i}$ and IIi estimations. ϕ_π declines to 1.28 and 1.44 with 4 lags, while ϕ_y declines to 1.09 and 1.23, respectively. These coefficient changes are somewhat larger than in the main estimation, and each remains statistically significant. The sum of the autoregressive coefficients does not statistically significantly change between subsamples, increasing to .91 in II$\widehat{i}$ and .87 in IIi.

Appendix C Additional Diagnostics

This section reports and discusses results of additional diagnostic analyses for model estimation not included in Section 6.

C.1 Non-monetary Structural Shocks

The monetary structural shocks for the NK and DSGE models are shown in Figure 3 of Section 6. Here, Figure C2 and Figure C3 plot the non-monetary structural shocks from the NK (η) and DSGE models (ε). In addition to the model parameters, the shock structure is highly variable between periods. The variance ratios of the shocks can be found in Table C9, and the autocorrelations are in Table C10.

In the VAR, the standard deviation of the shock is largley consistent across subsamples for both inflation and the output gap. For the NK model, the standard deviation of the output gap increases substaintially between periods I and II (from .06 to .29 in IIi and II$\hat{i}$). In contrast, the standard deviation of the inflation shock is varies by estimation strategy, increasing between I and IIi but decreasing when II$\hat{i}$ is used. The output gap shock's persistence is largely stable between the Great Moderation and post-Crisis, shifting from .79 to .83 and .87. Meanwhile, he inflation shock is highly persistent and near unity in period I but declines substantially in period II (to .37 in IIi and .65 in II$\hat{i}$), corresponding to the lower variance in the inflation shock.

For the DSGE model, the standard deviation of the productivity and wage markup shocks increase substantially (from .37 to .45 and .29 to .63, respectively). On the other hand, the spending shock decreases between samples I and II (from .40 to .30), while the risk premium shock is varies based on the estimation strategy used. Additionally, the only significant changes in non-monetary shock persistence are declines in the spending (.96 to .84 and .74), the price markup (.75 to .54 and .55), and wage markup shocks (.79 to .22 and .21)

C.2 Non-Monetary Impulse Responses

The main text focuses on the responses of the benchmark models to innovations in the monetary policy shock, which is common to all three models and directly relevant to the Taylor Rule. We do not include impulse responses to the other structural shocks here for two reasons. First, the non-monetary shocks are not easily compared across models, and second, sheer number of responses requires too much textual discussion. However, the full set of impulse responses is available upon request.

C.3 Output Gaps

Figure C1 plots the DSGE output gaps for all samples along with the CBO output gap for comparison. Unlike the CBO output gap, changes in model coefficients across subsamples influence the estimated DSGE output gap and cause discontinuities across subsamples. The Great Moderation DSGE and CBO gaps are positively correlated and have comparable levels until about 1992. After that the DSGE gap tells a drastically different story, remaining positive and large for nearly two decades. The magnitude of divergence is economically meaningful for monetary policy responses to output gaps in the Taylor Rule for all models. The divergence also may be a concern for construction and interpretation of the two gaps.

After the GFC, the post-Crisis (FF) (IIi), post-Crisis (Wu-Xia) (II$\hat{i}$), and CBO output gaps are highly correlated and follow a similar U-shaped path. However, the CBO gap returns to zero faster actually goes positive by 2019. In contrast, the post-Crisis (FF) and post-Crisis (Wu-Xia) gaps are still around -1 percent. This discrepancy has implications for the determination of optimal monetary (and fiscal) policy during the COVID-19 recession and recovery.

Subsample breaks in the DSGE output gap provide complementary evidence of structural breaks in the DSGE model coefficients. The results in Section 5 suggest that changes in long-run coefficients like the steady-state growth rate likely play an important role, but changes in coefficients associated with wage-price block and Taylor Rule may also contribute. Output gaps are notoriously difficult to estimate, as show by (Quast and Wolters, 2023), and the results in this appendix may reflect the impact of the omission of explicit UMP in the macro model equations. Either way, failure to allow for structural breaks in model coefficients appears to lead to bias in the estimated full-sample DSGE output gap for long periods.

Fig. C1 Estimated Output Gaps

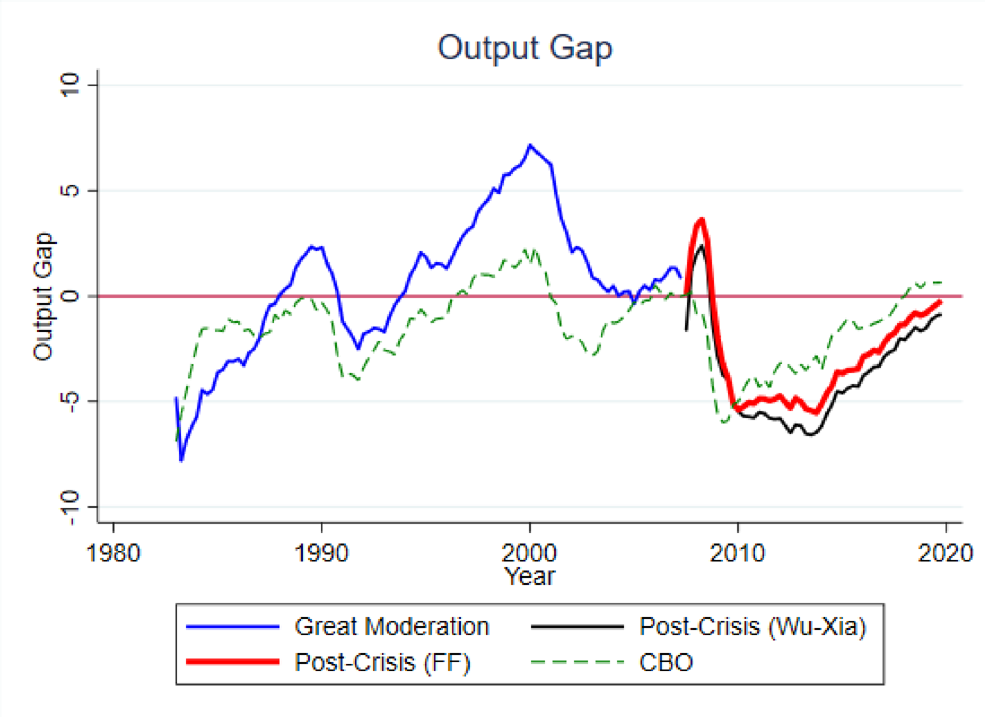


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Table 1 Summary of structural break literature

Paper	Sample	Break(s)	Taylor Rule Implications
Estrella and Fuhrer (2003)	1966-1997	early-1980s	backward looking models are more stable
Duffy and Engle-Warnick (2006)	1955-2003	1980	$\downarrow \phi_y$
Bernanke and Mihov (1998)	1965-1996	1979 & 1982	no policy variable captures monetary policy
Clarida et al. (2000)	1960-1996	1979	$\uparrow \phi_\pi, \uparrow \phi_y$
Smets and Wouters (2007)	1954-2007	early-1980s	$\uparrow \phi_\pi, \downarrow \phi_y$
Canova (2009)	1955-2002	1982	$\uparrow \phi_\pi$
Castelnuovo (2012)	1966-2007	1970s	Including M2 improves fit before 1970s
Ilbas (2012)	1966-2005	early-1980s	$\uparrow \phi_r, \uparrow \phi_\pi$
Fernández-Villaverde and Rubio-Ramírez (2007)	1955-2000	NA	Parameters tend to drift over time
Bunzel and Enders (2010)	1965-2007	early-1980s	$\uparrow \phi_r, \uparrow \phi_\pi$
Coibion and Gorodnichenko (2011)	1960-2002	early-1980s	$\uparrow \phi_\pi, \uparrow \phi_{\Delta y}, \downarrow \phi_y$
Sims and Zha (2006)	1959-2003	late-1970s, mid-1980s	$\downarrow$ variance of ϕ_π, ϕ_y
Primiceri (2005)	1953-2001	early-1980s	$\downarrow$ variance of $\phi_\pi, \phi_y, \uparrow \phi_\pi$
Mavroeidis (2010)	1961-2006	1979	TR cannot be accurately estimated after 1979
Carvalho et al (2021)	1960-2007	1979 & 1987	$\uparrow \phi_\pi, \downarrow \phi_Y, \pi^*$
Lakdawala (2016)	1965-2007	early-1980s	$\uparrow \phi_\pi, \downarrow \phi_Y$
Nikolsko-Rzhevskyy et al. (2019)	1965-2013	1979, 1987, 2001, 2007	$\uparrow \phi_\pi, \downarrow \phi_Y$
Dean and Schuh	1984-2019	2007	$\downarrow \phi_\pi$

Table 2 Taylor Rule estimates by subsample and model

Parameter Estimates						Changes	
	Parameter	Full	I	II$\hat{i}$	IIi	I-II$\hat{i}$	I-IIi
VAR	ϕ_π	1.74 [1.90, 2.50]	2.87 [2.27,3.07]	1.48 [.94,1.93]	1.29 [.44,2.16]	-1.39***	-1.59***
	ϕ_y	3.40 [3.32,3.68]	3.25 [2.93,3.27]	1.59 [1.36,1.73]	1.72 [.61,2.81]	-1.66***	-1.53***
	$\sum_{j=1}^{2} \rho_{i,j}$	.89 [.8,.98]	.89 [.74,1.04]	.88 [.68,1.08]	.82 [.58,1.04]	-.01	-.06
NK	ϕ_π	1.50 [1.07,1.86]	2.41 [2.30,2.55]	1.56 [1.20,1.92]	1.31 [1.00,1.57]	-.85***	-1.1***
	ϕ_y	.98 [.72,1.26]	.69 [.43,.99]	.94 [.70,1.20]	.64 [.38,.93]	.25	-.05
	ρ_i	.76 [.68,.84]	.62 [.54,.68]	.61 [.52,.69]	.78 [.68,.88]	-.01	.16*
DSGE	ϕ_π	1.78 [1.51,2.07]	1.97 [1.58,2.36]	1.44 [1.09,1.77]	1.48 [1.12,1.84]	-.53**	-.49**
	ϕ_y	.34 [.30,.38]	.09 [.02,.17]	.13 [.08,.18]	.17 [.10,.24]	.04	.08
	ρ_i	.83 [.78,.88]	.84 [.79,.88]	.83 [.76,.91]	.86 [.79,.93]	-.01	.02
	$\phi_{\Delta y}$	.33 [.29,.37]	.16 [.11,.21]	.16 [.11,.22]	.13 [.06,.19]	0	-.03

90% confidence interval in brackets; *** p<0.01, ** p<0.05, * p<0.1
Statistical significance of parameter differences across subsamples is computed using Wald tests.
Note: The VAR is estimated using OLS. The NK and DSGE are estimated with Bayesian methods using priors consistent with those used in previous estimations. In IIi, the VAR is estimated using the method in Aruoba et al (2022) while the NK and DSGE use the occasionally binding constraint method described in Cuba-Borda et al (2019) and Giovannini et al (2021).

Table 3 Structural Estimates from New Keynesian and Bayesian DSGE Model

	Parameter Output						Change	
	Parameter	Parameter Role	Full	I	IIi	II$\hat{i}$	I-IIi Change	I-II$\hat{i}$ Change
VAR	κ	Phillips Curve	.03 [-.03,.10]	-.03 [-.10,.05]	.13 [-.06,.21]	.15 [-.01,.32]	.16***	.18***
NK	ψ	IS Curve	-.1 [-.17,-.01]	-.03 [-.06,-.01]	-.20 [-.06,-.33]	-.19 [-.38,-.01]	-.17**	-.16
	κ	Phillips Curve	.04 [.02,.06]	.38 [.32,.42]	.05 [.02,.08]	.03 [.01,.04]	-.33***	-.35***
	β	Inflation feedback	.60 [.44,.75]	.91 [.87,.95]	.70 [.61,.80]	.82 [.61,.99]	-.21***	-.09
DSGE	$100(\beta^{-1}-1)$	Time Preference	.28 [.14,.41]	.17 [.06,.29]	.19 [.07,.28]	.17 [.07,.31]	.02	0
	$\bar{\pi}$	Steady State Inflation	.60 [.46,.74]	.68 [.56,.84]	.61 [.51,.71]	.65 [.50,.80]	-.07	-.03
	$\bar{\gamma}$	Steady State Growth	.37 [.32,.43]	.48 [.42,.53]	.22 [.17,.27]	.22 [.16,.28]	-.26***	-.26***
	$\bar{l}$	Steady State Hours	-2.28 [-3.8,-.79]	.79 [-1.36,2.46]	-4.00 [-5.16,-2.34]	-3.26 [-4.71,-1.97]	-4.79***	-4.05***
	ρ	Investment Adjustment	6.83 [6.5.29,8.44]	6.38 [4.14,8.66]	5.50 [3.18,7.38]	5.74 [3.76,7.38]	-.88	-.64
	σ_c	Risk Aversion	1.61 [1.18,2.03]	1.25 [.81,1.75]	1.10 [.74,1.43]	.94 [.72,1.14]	-.15	-.31
	λ	External Habit Degree	.65 [.57,.72]	.52 [.39,.67]	.63 [.50,.75]	.74 [.63,.85]	.11	.22
	ξ_w	Calvo: Wages	.78 [.72,.84]	.69 [.52,.89]	.82 [.74,.91]	.75 [.64,.88]	.13	.06
	σ_l	Frisch Elasticity	2.04 [1.37,2.65]	2.20 [1.16,3.37]	1.15 [.25,1.84]	.97 [.25,1.81]	-1.05*	-1.23*
	ξ_p	Calvo: Prices	.91 [.89,.93]	.81 [.74,.88]	.73 [.61,.86]	.72 [.60,.84]	-.08*	-.09
	ι_w	Wage Indexation	.38 [.21,.53]	.46 [.17,.74]	.37 [.16,.56]	.38 [.16,.58]	-.09	-.08
	ι_p	Price Indexation	.30 [.13,.50]	.35 [.12,.60]	.30 [.12,.48]	.32 [.14,.49]	-.05	-.03
	ψ	Capacity Utilization Cost	.78 [.66,.89]	.68 [.46,.88]	.74 [.59,.91]	.75 [.59,.91]	.06	.07
	Φ	Fixed Cost Share	1.60 [1.49,1.72]	1.53 [1.32,1.70]	1.36 [1.21,1.51]	1.39 [1.24,1.54]	-.17	-.14
	α	Capital Share	.20 [.17,.23]	.21 [.15,.26]	.13 [.09,.17]	.11 [.07,.15]	-.08*	-.1**
	r^*	Real Interest Rate	3.53	3.09	1.74	1.55	-1.35	-1.54

90% confidence interval in brackets; *** p<0.01, ** p<0.05, * p<0.1
Statistical significance of parameter differences across subsamples is computed using Wald tests.
Note: The VAR is estimated using OLS. The NK and DSGE are estimated with Bayesian methods using priors consistent with those used in previous estimations. In IIi, the VAR is estimated using the method in Aruoba et al (2022) while the NK and DSGE use the occasionally binding constraint method described in Cuba-Borda et al (2019) and Giovannini et al (2021).

Table A1 Estimation priors

Model	Parameter	Parameter Role	Prior Distribution	Prior Mean
NK	ψ	IS curve slope	Gamma	-.50 (.35)
	κ	Phillips Curve slope	Gamma	1.00 (2.00)
	β	Inflation Expectation feedback	Beta	.98 (.05)
	ρ_i	Monetary smoothing	Beta	.8 (.25)
	ϕ_π	Taylor Rule: Inflation	Normal	1.3 (.5)
	ϕ_y	Taylor Rule: Output	Beta	.5 (.25)

Table A2 Estimation priors

Model	Parameter	Parameter Role	Prior Distribution	Prior Mean
DSGE	$100(\beta^{-1}-1)$	Time Preference Rate	Gamma	.25 (.10)
	$\bar{\pi}$	Steady State Inflation	Gamma	.62 (.10)
	$\bar{\gamma}$	Steady State Growth Rate	Normal	.40 (.10)
	$\bar{l}$	Steady State Hours	Normal	.00 (2.00)
	ρ	Investment Adjustment Cost	Normal	4.00 (1.50)
	σ_c	Risk Aversion	Normal	1.50 (.37)
	λ	External Habit Degree	Beta	.70 (.10)
	ξ_w	Calvo Parameter: Wages	Beta	.50 (.10)
	σ_l	Frisch Elasticity	Normal	2.00 (.75)
	ξ_p	Calvo Parameter: Prices	Beta	.50 (.10)
	ι_w	Indexation to Past Wages	Beta	.50 (.15)
	ι_p	Indexation to Past Prices	Beta	.50 (.15)
	ψ	Capacity Utilization Cost	Beta	.50 (.15)
	Φ	Fixed Cost Share	Normal	1.25 (.12)
	α	Capital Share	Normal	.30 (.05)
	ρ	Monetary smoothing	Beta	.75 (.1)
	ϕ_π	Taylor Rule: Inflation	Normal	1.5 (.25)
	ϕ_y	Taylor Rule: Output	Normal	.12 (.05)
	$\phi_{\Delta y}$	Taylor Rule: Growth	Normal	.12 (.05)

Table B3 Taylor Rule estimates using the Taylor Rule mentioned in Carlstrom et al (2017)

	Parameter Estimates					Changes	
	Parameter	Full	I	II$\hat{i}$	IIi	I-II$\hat{i}$	I-IIi
VAR	ϕ_π	1.94 [.33, 1.82]	3.27 [2.05,4.50]	1.27 [.37,2.17]	1.42 [-.32,3.16]	-2***	-1.85*
	ϕ_y	1.25 [.75,2.42]	1.81 [1.28,2.34]	1.71 [.27,3.17]	1.06 [.45,1.66]	-.1	-.75***
	$\sum_{j=1}^{2} \rho_{i,j}$	.78 [.61,.95]	.88 [.71,1.05]	.87 [.71,1.03]	.84 [.66,1.02]	-.01	-.04
	ϕ_{tp}	2.32 [.91,98]	2.52 [1.85,3.19]	1.59 [.85,2.33]	1.98 [1.13,2.83]	-.93**	-.57

90% confidence interval in brackets; *** p<0.01, ** p<0.05, * p<0.1
Statistical significance of parameter differences across subsamples is computed using Wald tests.
Note: The VAR is estimated using OLS. In IIi, the VAR is estimated using the method in Aruoba et al (2022).

Table B4 Lag selection test for VAR

Lag	AIC	HQIC	BIC
1	2.67	2.76	2.91
2	2.42*	2.59*	2.84*
3	2.43	2.67	3.03
4	2.45	2.78	3.24
5	2.44	2.73	3.30
6	2.48	2.84	3.52

* indicates the information criterion's preference. Lag selection test is run over the entire sample.

Table B5 Taylor Rule estimates using real-time data

	Parameter Estimates					Changes	
	Parameter	Full	I	II$\hat{i}$	IIi	I-II$\hat{i}$	I-IIi
VAR	ϕ_π	1.30 [1.04, 1.56]	1.51 [.63,2.40]	.56 [-.24,.1.36]	.37 [.01,.73]	-.95**	-1.14***
	ϕ_y	1.16 [1.04,1.27]	2.20 [1.17,3.22]	.50 [-.29,1.30]	.33 [-.05,.70]	-1.70***	-1.87***
	$\sum_{j=1}^{2} \rho_{i,j}$	.95 [.83,1.07]	.93 [.79,1.07]	.91 [.74,1.08]	.85 [.68,1.03]	-.02	.08

90% confidence interval in brackets; *** p<0.01, ** p<0.05, * p<0.1
Statistical significance of parameter differences across subsamples is computed using Wald tests.
Note: The VAR is estimated using OLS. In IIi, the VAR is estimated using the method in Aruoba et al (2022).

Table B6 Taylor Rule estimates using 4 lags

Parameter Estimates						Changes	
	Parameter	Full	I	II$\hat{i}$	IIi	I-II$\hat{i}$	I-IIi
VAR	ϕ_π	1.99 [1.73,2.25]	2.54 [2.07,3.01]	1.28 [.37,2.19]	1.44 [1.02,1.86]	-1.26**	-1.1***
	ϕ_y	1.71 [1.59,1.81]	2.91 [2.39,3.45]	1.09 [.19,1.99]	1.23 [.79,1.67]	-1.82***	-1.68***
	$\sum_{j=1}^{4} \rho_{i,j}$	.94 [.79,1.09]	.84 [.65,1.03]	.91 [.69,.1.13]	.87 [.65,1.09]	.07	.03

90% confidence interval in brackets; *** p<0.01, ** p<0.05, * p<0.1
Statistical significance of parameter differences across subsamples is computed using Wald tests.
Note: The VAR is estimated using OLS. In IIi, the VAR is estimated using the method in Aruoba et al (2022).

Fig. C2 Structural shocks (η) from NK model

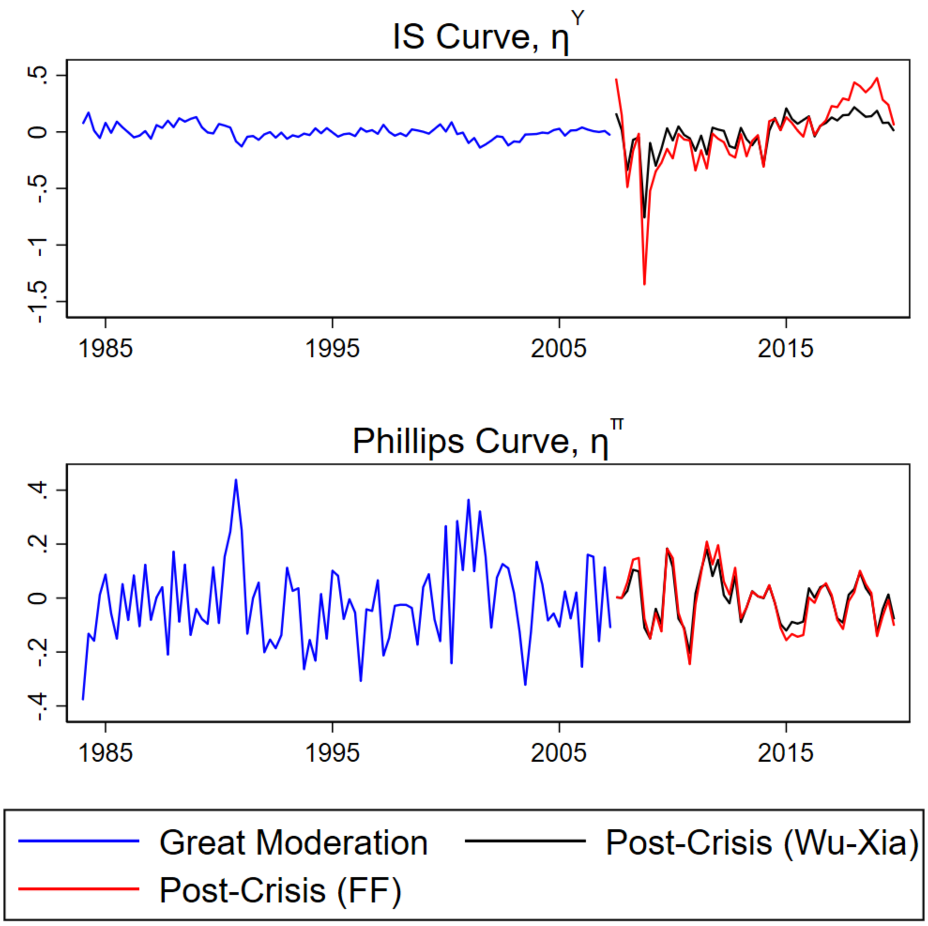

Fig. C3 Structural shocks (ε) from DSGE model

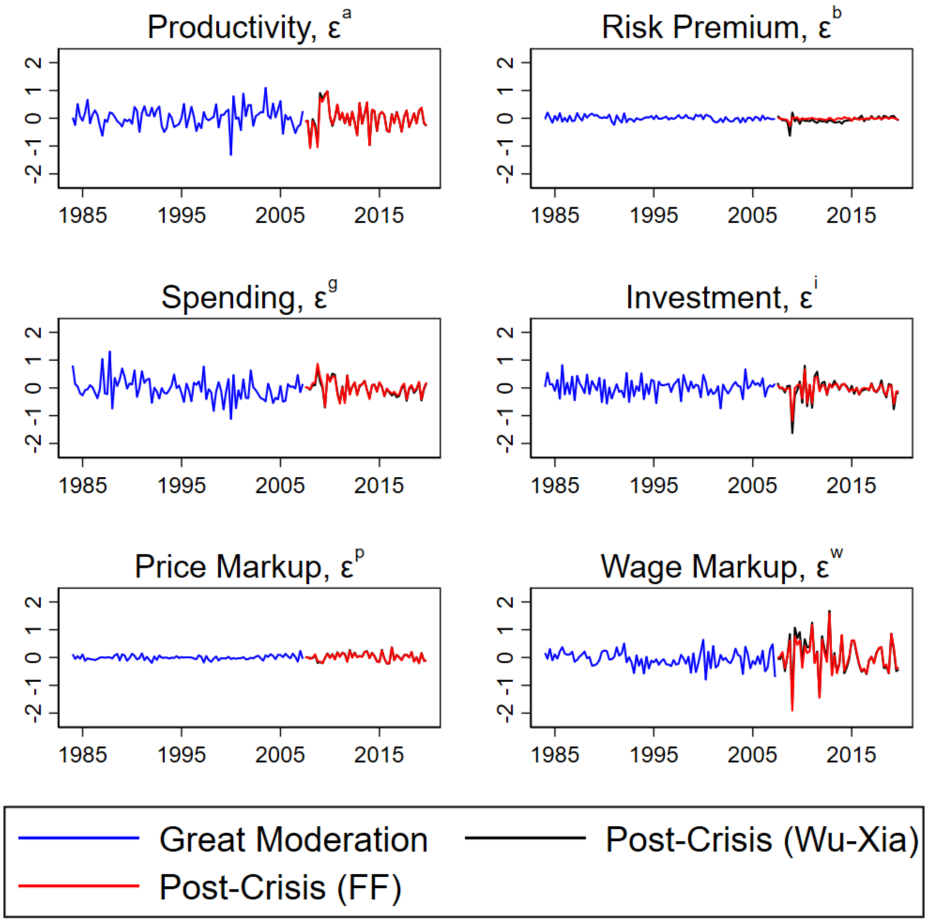

Fig. C4 Counterfactual responses to a 100bp monetary shock by model and sample

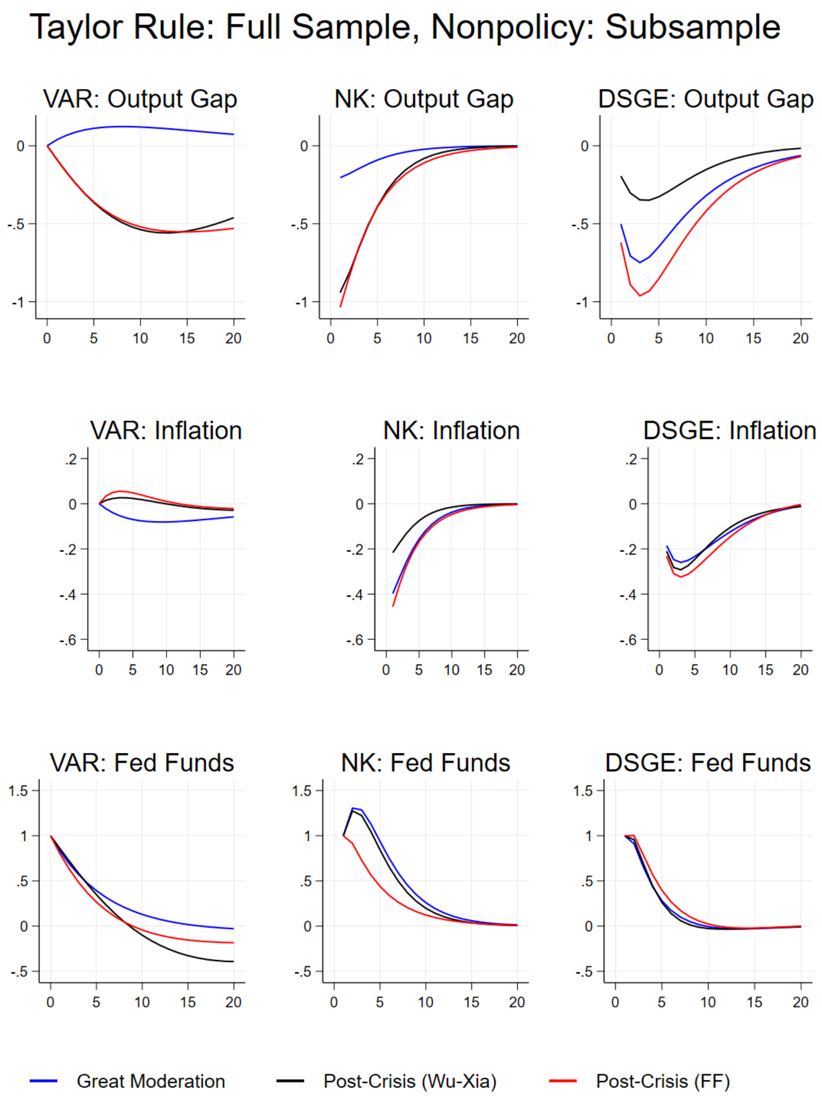

Table C7 Taylor Rule estimates by subsample and model

Parameter Estimates		
	Parameter	1960-1979
VAR	ϕ_π	1.52 [1.03,2.01]
	ϕ_y	1.07 [.85,1.29]
	$\sum_{j=1}^{2} \rho_{i,t-j}$	.80 [.63,.97]
NK	ϕ_π	1.27 [1.00,1.47]
	ϕ_y	.85 [.62, 1.06]
	ρ_i	.63 [.50, .74]
DSGE	ϕ_π	1.48 [1.26, 1.69]
	ϕ_y	.15 [.09, .20]
	ρ_i	.80 [.74, .87]
	$\phi_{\Delta y}$	.17 [.13, .22]

Note: 90% confidence interval in brackets. Estimations use the same strategy and priors as in the main text

Table C8 Structural Estimates from New Keynesian and Bayesian DSGE Model

	Parameter	Parameter Role	1960-1979
VAR	κ	Phillips Curve	-.01 (.04)
NK	ψ	IS Curve	-.07 [-.12,-.01]
	κ	Phillips Curve	.02 [.01,.03]
	β	Inflation feedback	.69 [.53,.87]
DSGE	$100(\beta^{-1}-1)$	Time Preference	.18 [.07,.30]
	$\bar{\pi}$	Steady State Inflation	.70 [.51,.86]
	$\bar{\gamma}$	Steady State Growth	.27 [.17,.38]
	$\bar{l}$	Steady State Hours	2.72 [1.28,4.33]
	ρ	Investment Adjustment	4.70 [3.11,6.19]
	σ_c	Risk Aversion	1.64 [1.25,1.98]
	λ	External Habit Degree	.67 [.60,.75]
	ξ_w	Calvo: Wages	.75 [.67,.83]
	σ_l	Frisch Elasticity	1.97 [1.09,3.09]
	ξ_p	Calvo: Prices	.55 [.50,.60]
	ι_w	Wage Indexation	.48 [.28,.67]
	ι_p	Price Indexation	.37 [.16,.62]
	ψ	Capacity Utilization Cost	.28 [.12,.42]
	Φ	Fixed Cost Share	1.55 [1.40,1.68]
	α	Capital Share	.24 [.19,.28]
	r^*	Real Interest Rate	2.50

Note: Estimations use the same strategy and priors as in the main text

Table C9 Structural shock standard deviation

	Parameter	Role	I	IIi	II$\widehat{i}$
VAR	σ_t^y	Output Gap	.47	.54	.53
	σ_t^π	Inflation	.20	.18	.18
	σ_t^i	Monetary	.47	.27	.33
NK	η_t^y	Output Gap	.06 [.04, .09]	.29 [.21, .36]	.33 [.12, .53]
	η_t^π	Inflation	.18 [.14, .22]	.23 [.18, .28]	.10 [.06. .13]
	η_t^m	Monetary	.40 [.33, .46]	.32 [.31, .33]	.35 [.28, .42]
DSGE	ϵ_t^a	Productivity	.37 [.31, .43]	.45 [.38, .53]	.45 [.37, .53]
	ϵ_t^b	Risk Premium	.09 [.05, .18]	.06 [.03, .08]	.14 [.10, .17]
	ϵ_t^g	Spending	.40 [.34, .46]	.30 [.25, .35]	.30 [.24, .35]
	ϵ_t^q	Investment	.31 [.23, .41]	.29 [.18, .41]	.38 [.20, .53]
	ϵ_t^m	Monetary	.12 [.10, .14]	.10 [.06, .13]	.34 [.29, .40]
	ϵ_t^p	Price Markup	.08 [.06, .10]	.15 [.11, .18]	.15 [.11, .18]
	ϵ_t^w	Wage Markup	.29 [.22, .26]	.63 [.51, .73]	.64 [.52, .76]

Table C10 Structural shock persistence

	Parameter	Role	I	IIi	II$\widehat{i}$
NK	ρ_y	Output Gap	.79 [.74, .85]	.83 [.78, .88]	.87 [.78, .99]
	ρ_π	Inflation	.99 [.99, 1]	.37 [.19, .57]	.65 [.42, .84]
	ρ^m	Monetary	.60 [.41, .71]	.22 [.03, .37]	.51 [.26, .81]
DSGE	ρ_a	Productivity	.92 [.87, .98]	.81 [.73, .94]	.85 [.78, .91]
	ρ_b	Risk Premium	.74 [.22, .93]	.90 [.85, .95]	.78 [.72, .86]
	ρ_g	Spending	.96 [.93, .99]	.84 [.74, .94]	.74 [.63, .93]
	ρ_q	Investment	.70 [.55, .84]	.75 [.55, .93]	.58 [.31, .82]
	ρ_m	Monetary	.31 [.13, .51]	.62 [.43, .84]	.51 [.33, .69]
	ρ_p	Price Markup	.75 [.55, .93]	.54 [.22, .82]	.55 [.31, .81]
	ρ_w	Wage Markup	.79 [.50, .97]	.22 [.06, .37]	.21 [.05, .37]

Note: The VAR's shock persistence is the autoregressive parameter from each equation in the model.