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ARTICLE

New evidence on US monetary policy activism and the Taylor rule

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Abstract
We provide new evidence about US monetary policy using a model that: (i) estimates time-varying monetary policy weights without relying on stylized theoretical assumptions; (ii) allows for endogenous breakdowns in the relationship between interest rates, inflation, and output; and (iii) generates a unique measure of monetary policy activism that accounts for economic instability. The joint incorporation of endogenous time-varying uncertainty about the monetary policy parameters and the stability of the relationship between interest rates, inflation, and output materially reduces the probability of determinate monetary policy. The average probability of determinacy over the period post-1982 to 1997 is below 60% (hence well below seminal estimates of determinacy probabilities that are close to unity). Post-1990, the average probability of determinacy is 75%, falling to approximately 60% when we allow for typical levels of trend inflation.

Keywords: Monetary policy; interest rates; inflation; Taylor rule; determinacy

JEL classifications: E52; E40; C11; C32

1. Introduction

There is no agreement on what the Taylor rule weights on inflation and the output gap should be, except with respect to their signs. The optimal weights would respond not only to changes in preferences of policymakers but also to changes in the structure of the economy and the channels of monetary policy transmission (Bernanke, 2015).

There is considerable uncertainty about the appropriateness of recent monetary policy in the US, including the extent to which monetary policy exacerbated the Great Recession. Taylor (2007, 2012) argues that US interest rates were overly accommodative following the 2001 slump, thereby contributing to unsustainable house price appreciation and a subsequent economic collapse. This perspective is opposed by Bernanke (2010, 2015) who highlights the uncertainty associated with the parameters underpinning the extent to which monetary policy should target inflation or output gaps. There are also a number of competing explanations about monetary policy and economic conditions during other periods of interest, such as the Great Inflation of the 1970s. These include passive monetary policy (Clarida, et al. 2000; Lubik and Schorfheide, 2004), lack of information and data for appropriate monetary policy (Orphanides, 2001), and the adverse effects of major disturbances (Sims and Zha, 2006).

We contribute to this discussion by formulating and estimating a model of the relationship between US interest rates, inflation, and output over the period 1955 to 2019 that: 1) identifies and estimates the monetary policy weights on inflation and output at every time point; 2) generates a unique measure of the exact probability of active monetary policy at every time point that accounts for data, parameter, and model uncertainty; and (3) allows us to examine the time-varying ramifications of monetary policy activism and determinacy in the US economy after endogenously accounting for economic instability. Our findings are important for understanding the post-WW2 evolution of US monetary policy, examining the relationship between monetary policy and economic growth, and designing appropriate monetary policy rules.

A unique feature of our approach is that we decompose monetary policy determinacy into the components associated with: (i) the probability of satisfying the Blanchard–Kahn conditions (Blanchard and Kahn, 1980); and (ii) the probability of a common stochastic relationship between interest rates, inflation, and output. In this respect, a common feature in the literature on modeling US monetary policy and estimating the probability of monetary policy determinacy and activism is that estimation of equations such as

$$i_t = (1-\rho)i_t^* + \rho i_{t-1} + e_t$$
$$i_t^* = \beta_\pi \pi_t + \beta_y y_t$$

inherently assume the presence of a permanent common stochastic relationship between interest rates (i_t), inflation (π_t), and the output gap (y_t).[1] We show that assuming a permanent common relationship between interest rates, inflation, and output (irrespective of whether the model parameters (e.g. β_π, β_y) are constant or allowed to vary over time) contaminates estimated monetary policy parameters. This has important policy ramifications because it loads the probabilities of monetary policy determinacy and activism entirely on the stance of the monetary authority (viz. it assumes that the monetary authority could have altered conditions had it adopted a different stance).

By relaxing the restriction of a permanent common stochastic relationship between interest rates, inflation, and output, we show that much of the indeterminacy in monetary policy is attributable to instability in the relationship between interest rates, inflation, and output, which cannot be controlled by simply changing the stance of monetary policy. Moreover, relaxing the aforementioned restriction implies a far more aggressive targeting of inflation than is typically estimated in the literature.

In the case of the Great Inflation of the 1970s, our findings largely reject the notion that a different monetary policy stance would have materially altered realized inflation. In particular, we show that estimates of monetary policy passivity and indeterminacy in the 1970s are largely attributable to instability in the rank of the relationship between interest rates, inflation, and output rather than due to either the probability of actively targeting inflation or the probability of satisfying the Blanchard–Kahn conditions. Moreover, the shift from indeterminate to determinate monetary policy during the Volcker regime is shown to heavily reflect improved stability in the relationship between inflation, interest rates, and output gap (as opposed to simply a change in the stance of monetary policy).

Historically, both reduced-form and structural approaches have been used to estimate monetary policy weights and monetary policy activism probabilities (Clarida, et al. 2000; Lubik and Schorfheide, 2004; Cogley and Sargent, 2005; Boivin, 2006; Sims and Zha, 2006; Fernández-Villaverde and Rubio-Ramirez, 2008). In the case of reduced-form models, their key benefit is that estimates are not contingent on a range of structural assumptions. A drawback, however, is the presence of confounding effects and biases stemming from alternative linear relationships that may be present in the data; for example, the relationship between interest rates, inflation, and output is typically explained by linear relationships governing monetary policy rules, the Phillips Curve, and an Euler equation. In a reduced-form framework (e.g. a regression of interest rates on inflation and the output gap), these alternative relationships effectively yield monetary policy parameters that are potentially hybrids of a range of linear relationships present in the data. While structural models explicitly account for the aforementioned linear relationships, they are contingent on strict theoretical assumptions that often only weakly reflect the underlying data. Moreover, in both the reduced-form and structural cases, model estimates are contingent on the strong assumption that common stochastic relationships cannot break down (e.g. a Taylor-type rule that is always active).

Our approach incorporates features of both the reduced-form and structural approaches and is based on the estimation of a novel Vector Error Correction Model (VECM) that allows for time-variation in both the parameters and the number of common stochastic relationships present in the data. Importantly, we also impose restrictions that allow for the unique identification of the parameters relating interest rates to inflation and output without the need to explicitly identify separate equations for the alternative linear relationships that may be present in the data (such as a Taylor-rule equation, a Phillips Curve equation and an equation based on Euler conditions). Estimation is undertaken using the methodology presented in Chua and Tsiaplias (2018), which allows for the estimation of models involving endogenous time-variation in both the cointegrating rank and parameters of a multivariate system.

A key feature of the approach adopted in this paper is that, in contrast to the existing literature, we estimate the time-varying monetary policy weights without imposing the restriction that there is a common stochastic relationship between interest rates, inflation, and output at all time periods. Instead, we endogenously account for economic instability that can generate a breakdown in the relationship between interest rates, inflation, and output. Econometrically, the typical restriction of a common stochastic relationship between interest rates, inflation, and output at all times can be characterized as a restriction on the rank of the system governing the relationship between the three key variables; the restriction is consistent with a system where it is assumed that the rank of the system is always equal to unity (viz. rank = 1 such that a single, common relationship always exists between interest rates, inflation, and output). To see why such a restriction is onerous, consider the situation where the number of common stochastic relationships in period t is zero. During this period, there is a breakdown in the relationship between the three key variables such that the parameters that govern whether monetary policy is active or determinate are not identified.

We show that the traditional imposition of the restriction of a permanent stochastic relationship between interest rates, inflation, and output is not only an econometric limitation but has significant economic ramifications for the monetary policy weights attached to inflation and output. In particular, relaxing the restriction of a permanent common stochastic relationship between interest rates, inflation, and output persistently lowers the probability of monetary policy determinacy.

There are three key benefits to removing the aforementioned restriction of a permanent, common stochastic relationship between interest rates, inflation, and output. First, the monetary policy weights are estimated in a more realistic setting that endogenously allows for economic instability that can lead to a breakdown in the relationship between interest rates, inflation, and output. This is achieved by allowing for the absence of any common stochastic relationship between the key variables (i.e. allowing for a time-varying rank of zero rather than permanently imposing a rank of unity). In this respect, evidence is provided of significant instability and persistent breakdowns in the relationship between interest rates, inflation, and output.

Second, we do not need to a priori impose the number of linear relationships present in the data (for example, assuming that monetary policy, Phillips curve, and Euler equations must always be present). Instead, we allow for up to three common stochastic relationships thereby enabling the identification of *all* possible long-run linear relationships that may be present among interest rates, inflation, and output. As such, we do not impose the presence of a permanent Taylor-type rule but allow for its existence if it can be uniquely identified in the data.

Third, we are able to construct new measures of monetary policy activism and determinacy which eliminate key limitations of existing measures. In previous research, the probability of a common stochastic relationship between interest rates, inflation, and output is always implicitly set to unity when determining the probability of determinacy (e.g.: Clarida, et al. 2000; Lubik and Schorfheide, 2004; Cogley and Sargent, 2005). A key issue with this approach is that it fails to account for the stability of the relationship between interest rates, inflation, and output. It therefore allows for statistically anomalous situations such as a probability of determinacy being close to unity even during periods when there is a high probability of a breakdown in the relationship between the key variables of interest.

Another benefit of our approach is that we uniquely identify the simultaneous relationships between interest rates, inflation, and output. This is non-trivial, with time-variation in the parameters typically resulting in estimated relationships that are not uniquely identified. To ensure that the parameters are uniquely identified, we impose the restriction that, at every time point, the parameters corresponding to each common stochastic linear relationship are orthogonal to the parameters of any other common stochastic relationship. This is important as we primarily observe shifts between zero and one common stochastic relationship between interest rates, inflation, and output. The former indicates a breakdown in the relationship between the three key variables, whereas the latter relationship is observationally equivalent to a Taylor-type rule where interest rates respond to inflation and output, but with weights associated with inflation and output targeting that vary over time.

The resulting methodology yields robust estimates of the continuous evolution of the monetary policy weights over the past 65 years. This is achieved in a model-consistent manner without the confounding effects and biases stemming from possible alternative linear relationships in the data or the need to impose strict theoretical assumptions that only weakly reflect the underlying data. In this respect, although large-scale DSGE models allow for multiple types of shocks, they are less likely to be effective in exercises such as measuring weights over extended periods of time. In practice, we do not know the true model of the economy, and the validity of any postulated economic structure will likely decline over long periods of time such as the period we consider here. In contrast, the setting we propose lends itself well to identifying the relationship between interest rates, inflation, and output over an extended period of time, allowing for the unique identification of monetary policy weights whilst avoiding the need to make assumptions about other aspects of the economy.

We note that there are other papers that also track the evolution of the monetary policy weights, usually using either time-varying models (e.g.: Primiceri, 2005; Boivin, 2006) or regime-switching models (e.g.: Sims and Zha, 2006; Davig and Leeper, 2007; Liu, et al. 2011; Bianchi, 2013; Davig and Doh, 2014). Primiceri (2005), for example, identifies the time-varying impact of monetary policy using a time-varying parameter (TVP) VAR. Bianchi (2013) estimates a model that examines the evolution of the weights using a Markov regime-switching approach that identifies dovish and hawkish periods. A key difference in this paper relative to the TVP or Markov-switching literature is that the time-variation is not limited to the VAR parameters, but also extends to the number of common stochastic relationships present between the variables. Moreover, the parameters in every possible common stochastic relationship are uniquely identified. These differences are significant, both econometrically and in terms of the resulting monetary policy weights.[2]

The remainder of the paper is structured as follows. Section 2 describes the econometric framework used to estimate the monetary policy weights. The results are discussed in Section 3, with Section 4 concluding.

2. Modelling the time-varying relationship between interest rates, inflation, and output

To estimate a fully time-varying system that allows us to quantify the pass-through from inflation and output to nominal short-term interest rates, a generalized error-correction model is formulated. We use the term 'generalised' to denote that the specification is similar to a VECM but

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significantly broader, with the model handling both stationary and non-stationary series in a time-varying parameter context. The model also allows for the unique and consistent identification of the time-varying pass-through parameters from inflation and output to interest rates.

The resulting parameter estimates do not rely on any a priori assumptions regarding the stationarity of the data (or even the presence of a cointegrating relationship), and the model encompasses time-varying structural elements that allow for periods when the data behave as I(0) (viz. integrated of order zero); I(1) and non-cointegrated; or I(1) with a cointegrating relationship. The literature in this space has typically relied on assumptions about whether the data are permanently I(0) or I(1), in addition to assumptions regarding the permanence of cointegrating relationships. We relax the need for such assumptions while still estimating models that take on forms that are consistent with beliefs about whether the data are I(0) or I(1). We note, however, that there are also alternative models that can be estimated, such as those based on fractional integration (although there is still the consideration of appropriately dealing with time-varying parameters and instability in the fractional context). Importantly, the model is able to account for periods where variables that are usually stationary behave more like non-stationary variables for a given period of time; during such periods, a policymaker will exhibit a greater level of uncertainty regarding the relationship between interest rates, inflation, and output. This uncertainty is accounted for in deriving the monetary policy weights.

The data generating process for the relationship between interest rates (i_{t+1}), inflation (π_{t+1}), and the output gap (y_{t+1}) at time $t+1$ is given by

$$\Delta x_{t+1} = \left[\Delta i_{t+1}, \Delta\pi_{t+1}, \Delta y_{t+1}\right] = c_{t+1} + \Delta x_t B_t + x_t \Theta_t + \epsilon_{t+1} \tag{1}$$

$$\epsilon_{t+1} \sim MVN(0, \Sigma_{t+1}) \tag{2}$$

where x_t is an $n=3$ dimensional row vector,[3] c_{t+1} is an $(1 \times n)$ dimensional time-varying intercept, B_t is an $(n \times n)$ coefficient matrix governing time-varying autoregressive dynamics, Θ_t is an $(n \times n)$ multiplier matrix for x_t, and ε_{t+1} is a $(1 \times n)$ independent multivariate Gaussian process with positive definite covariance matrix Σ_{t+1}.

Although all of the model's parameters are time-varying, the nature of the time-variation differs for Θ_t. The rank of Θ_t, governing the number of contemporaneous relationships present in the system during period t, depends on a discrete but unobserved Markovian regime process S_t that is discussed in Section 2.1. The Markovian regime S_t also governs the time-variation in c_{t+1}, B_t and Σ_t such that $c_{t+1} = c_{S_{t+1}}$, $B_{t+1} = B_{S_{t+1}}$ and $\Sigma_{t+1} = \Sigma_{S_{t+1}}$. The multiplier parameters in Θ_t, however, are subject to their own time-varying dynamics, which are independent of S_t (hence independent of the rank of Θ_t). The time-varying dynamics of the parameters in Θ_t, which determine the pass-through from inflation and output to interest rates, are discussed in Sections 2.2 and 2.3.

Consequently, there are two independent sources of variation in the multiplier matrix used to determine the contemporaneous relationships in x_t: 1) time-variation in the number of contemporaneous relationships between the variables, and 2) time-variation in the multiplier parameters, allowing the dynamics of the contemporaneous relationship(s) to be time-varying. The first source of time-variation is based on discrete Markovian regimes and allows for simultaneous relationships between the variables (such as Taylor-type rules involving interest rates that respond to inflation and output), in addition to periods when the relationship between the variables in x_t breaks down. The second source of variation allows for time-variation (independent of the prevailing Markovian regime) in the underlying parameters of any simultaneous relationship in the data (such as time-variation in the weights on inflation and output stemming from a Taylor-type rule).

To see how the model can be used to identify the contemporaneous relationship between the variables in x_t, note that we can re-write equation (1) to show that the model produces a matrix of time-varying simultaneous relationships that accounts for both a time-varying intercept, and the leads and lags of x_t.

$$x_t\Theta_t = x_t B_t^*(L) - c_{t+1} + \epsilon_t^* \tag{3}$$

where $B_t^*(L) = \left[IL^{-1} \ -(I+B_t)L^0 \ B_tL^1 \right]$ is a polynomial in the lag operator L, resulting in residuals ϵ_t^* that are both stationary and serially uncorrelated. We discuss how equation (3) can be used to identify the monetary policy weights associated with Taylor-type rules in the next two sub-sections.

2.1. Allowing for time-variation in the rank of the multiplier matrix

To generate time-variation in the rank of Θ_t, a singular value decomposition is adopted such that $\Theta_t = U_t\Lambda_t V_t' = U_t D_t$, where U_t and V_t are orthonormal matrices of dimension n and Λ_t is a diagonal matrix of dimension n (Kleibergen and van Dijk, 1998; Kleibergen and Paap, 2002).

Following Chua and Tsiaplias (2018), the rank of Θ_t can be expressed as the following function of the idempotent matrix $I(S_t)$

$$\Theta_t = U_t I(S_t) I(S_t) D_t \tag{4}$$

where $I(S_t)$ is a diagonal matrix with kth diagonal element

$$I(S_t)_{kk} = (1 - s_{1t}) \sum_{j=k+1}^{n+1} s_{jt}, \quad k = 1 \text{ to } n. \tag{5}$$

The indicator variable s_{jt}, $j = 1$ to $n+1$, is set to unity if $S_t = j$ and to zero if $S_t \neq j$.

Given (5), when s_{jt} is equal to 1 then the rank of Θ_t is $(j-1)$. The rank of Θ_t is therefore uniquely determined by the state variable S_t. The probability of a move from state i to state j (being also the probability of a change in the rank of Θ_t) is given by

$$p_{ij} = \Pr(S_t = j | S_{t-1} = i), \qquad \sum_{j=1}^{n+1} p_{ij} = 1. \tag{6}$$

The transition process (6) can be represented in matrix form as[4]

$$P = \begin{bmatrix} p_{11} & p_{12} & p_{13} & p_{14} \\ p_{21} & p_{22} & p_{23} & p_{24} \\ p_{31} & p_{32} & p_{33} & p_{34} \\ p_{41} & p_{42} & p_{43} & p_{44} \end{bmatrix}. \tag{7}$$

It can be shown that Θ_t can be expressed as the product of two matrices, β_t and α_t, pursuant to

$$\beta_t = U_t I(S_t), \tag{8}$$

$$\alpha_t = I(S_t)\Lambda_t V_t' = I(S_t) D_t. \tag{9}$$

Consequently, in the case where $S_t = (r+1)$ and $r < n$,[5] the rank of Θ_t is r and the model produces β_t and α_t that satisfy

$$\beta_t = U_{t,1:r},$$

$$\alpha_t = \Lambda_{t,1:r,1:r} V_{t,1:r}' = \left(D_{t,1:r}'\right)',$$

where $U_{t,1:r}$ and $V_{t,1:r}$ are the first r columns of U_t and V_t, respectively, and $\Lambda_{t,1:r,1:r}$ is the upper left $r \times r$ submatrix of Λ_t.

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When $S_t = 2$ (implying $r = 1$), nominal interest rates, inflation, and output share a common stochastic relationship. This restriction is imposed in the existing literature prior to estimation, thereby not allowing for the possibility of any other state other than $S_t = 2$. In our setting, if $p_{i2} = 1$ for any i then S_t will be equal to 2 for any t, thereby ensuring the system is always defined by a single common stochastic relationship between the variables in x_t. Given the aforementioned condition, the multivariate system for x_t can be expressed as

$$\left[i_t - \beta_{\pi,t}\pi_t - \beta_{y,t}y_t\right]\alpha_t = x_t B_t^*(L) - c_{t+1} + \epsilon_t^* \tag{10}$$

where interest rates i_t respond to inflation via $\beta_{\pi,t}$ and to the output gap via $\beta_{y,t}$. Equation (10) holds if the rank of the system is equal to unity (according to $S_t = 2$), whereby the response of interest rates to inflation and output is characterized by a single equation.

2.2. Identifying the time-varying parameters in the multiplier matrix

To obtain time-consistent estimates of the critical pass-through parameters $\beta_{\pi,t}$, $\beta_{y,t}$, it is necessary to ensure that the parameters in the multiplier matrix are uniquely identified in each state of the system. This ensures that the multipliers in a given state at time t are never misidentified with the time-varying parameters in any other state.[6] The elimination of any misidentification must hold for every time t in order for the pass-through parameters associated with a given state to be continuously identified.

We achieve identification by imposing a unique orthogonality condition on the state-dependent multiplier parameters at every time point. Subject to this identification, it follows that if the system is always in either the first or second state, then there will be—at most—a single relationship between the levels of interest rates, inflation, and output in the data. If the reader is prepared to assume that the relationship estimated when $S_t = 2$ reflects a Taylor-type rule in even one particular period then the orthogonality condition implies that this interpretation will hold for *every* period. Consequently, if the system is never in the third or fourth states, there will be no other orthogonal interpretation of the relationship between the levels of interest rates, inflation, and output that is consistent with the data. To our knowledge, there is no competing model that is able to satisfy this restriction, thereby allowing us to consistently infer the entire time path of the monetary policy weights from the 1950s onwards, in addition to obtaining consistent data-driven estimates of the probability of active monetary policy at each time point.

We note that, even if the probabilities tend to favor $S_t = 2$, our approach will not necessarily generate the same results as previous research (which assumes that $S_t = 2$ with probability 1 for any t). We examine the empirical implications of imposing the latter restriction, including the size of the associated bias stemming from the assumption of a permanent common relationship, in Section 3.2.

To estimate the time-varying monetary policy weights subject to our orthogonality restriction, we first follow Koop et al. (2011) in introducing a parameter expansion that simplifies estimation by allowing us to estimate latent matrices D_t^* and U_t^* (hence $\alpha_t^* = I(S_t)D_t^*$ and $\beta_t^* = U_t^* I(S_t)$) in place of D_t and U_t (which are used to construct the multiplier matrix $\Theta_t = U_t I(S_t) I(S_t) D_t$ defined in equation 4).[7]

An issue, however, with the parameter expansion proposed by Koop et al. (2011) is that it does not eliminate the possibility that the state-dependent multiplier parameters change every time the state of the system changes (e.g. going from $S_t = 2$ to $S_t = 3$ produces a new set of multiplier parameters for both the first and second contemporaneous relationships). This is important as it prevents the unique identification of the monetary policy parameters. To correct for this, we follow Chua and Tsiaplias (2018) in adopting a parameter expansion that is unique in that it satisfies the restrictions imposed on the multiplier's parameter space without introducing a potential 'discontinuity' in the state-dependent multiplier parameters each time the state of the system changes.

An example of such a discontinuity is that the parameters in the first column of Θ_t can change when the prevailing state changes. In the presence of such discontinuity, the parameters associated with a given state can no longer be given a time-consistent interpretation; there is no longer a vector describing the relationship between the variables (e.g. a "Taylor-type" rule) that is legitimately "activated" or "deactivated" at each time period in accordance with the state of the system, nor can the probability of active monetary policy be consistently estimated over different points in time.

The discontinuity is eliminated by imposing orthogonality restrictions on both U_t^* and D_t^* such that the parameters in each column of Θ_t are uniquely identified by reference to S_t. By eliminating the discontinuity, we can estimate the probability of active monetary policy at every time point. While this issue would be unimportant in a purely reduced-form exercise (for example, for forecasting interest rates), it is critical when seeking to consistently interpret the time path of the monetary policy weights. Technical details regarding the imposition of the orthogonality restriction are left to Appendix A.

2.3. Specifying the dynamics of the parameters in the multiplier matrix

To operationalize the model, we specify the dynamics of the individual parameters in the U_t^* and D_t^* matrices (where the individual parameters are denoted $u_t^* = vec\left(U_t^*\right)$ and $d_t^* = vec\left(D_t^*\right)$ respectively) following Koop et al. (2011). By specifying autocorrelated dynamics, we are able to obtain estimates of the pass-through parameters even where the probability of $S_t = 2$ is small. We do this in a flexible manner by modeling the individual unrestricted elements in u_t^* as AR(1) processes

$$u_t^* = \rho u_{t-1}^* + \eta_t, \tag{11}$$

$$\eta_t \sim N\left(0, I_{n^*}\right), \tag{12}$$

where $u_0^* \sim N\left(0, I_{n^*}\frac{1}{1-\rho^2}\right)$ and $|\rho| < 1$.

The individual unrestricted elements in d_t^* that are used to construct α_t^* are modeled as flexible independent random walk processes such that

$$d_t^* = d_{t-1}^* + \zeta_t, \tag{13}$$

$$\zeta_t \sim N\left(0, Q\right), \tag{14}$$

where Q is a $(n^* \times n^*)$ diagonal matrix with elements $\left(\sigma_1^2, \ldots, \sigma_{n^*}^2\right)$ and $d_0^* \sim N\left(0, \sigma_d \times I_{n^*}\right)$. Note that the distinction between the autoregressive dynamics in (11) and the random walk dynamics in (13) is not arbitrary and is motivated by the necessity to rule out undesirable properties in the multiplier coefficients (see, further, Koop et al., 2011).

The specifications in equations (11) and (13) are important because they imply an inter-temporal relationship (in effect, a smoothness condition) between the state-dependent multiplier parameters. This allows us to form estimates of the pass-through even when the probability of a Taylor-type rule is small. Relatedly, the entire vector of parameters in $u_t^* = vec(U_t^*)$ and $d_t^* = vec\left(D_t^*\right)$ for $t = 2, \ldots, T$ can be efficiently estimated conditional on the autoregressive dynamics in equations (11) and (13), the orthogonality conditions described in Appendix A, the state vector S_t, and the observed data (see Steps 3 and 5 in Appendix B). Even if the probability of $S_t = 2$ is small, the estimates of the monetary policy pass-through parameters $\beta_{\pi,t}$ $\beta_{y,t}$ can be derived conditional on $S_t = 2$, the data x_t, the entire time path of $\beta_{\pi,t}$ and $\beta_{y,t}$, and the other parameters in d_t^* and u_t^* (which are related to $\beta_{\pi,t}$ and $\beta_{y,t}$ via the orthogonality conditions).

When coupled with the orthogonality condition, the inter-temporal relationship between the state-dependent multiplier parameters ensures that the resulting state-dependent parameters are uniquely identified both inter-temporally and across alternative regimes, thereby allowing for

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the identification and estimation of the time-varying monetary policy weights associated with inflation and output in a time-consistent manner, but without imposing any stylized theoretical assumptions (such as assumptions on the rationality of expectations, the rigidity of prices, or the stickiness of consumption).

2.4. Interest rate smoothing

In terms of the linear stochastic relationship governing monetary policy, a Taylor-type rule assumes a general form for the nominal federal funds target rate of

$$i_t^* = \beta_\pi \pi_t + \beta_y y_t. \tag{15}$$

More generally, the above process can be written with the inclusion of an intercept. We assume demeaned variables and omit the intercept.[8]

To better fit actual interest rates i_t, interest rate smoothing is often introduced. For example, actual interest rates may follow

$$i_t = (1 - \rho) i_t^* + \rho i_{t-1} + e_t \tag{16}$$

where e_t is a shock and $\rho \neq 0$ reflects the presence of (typically positive) autocorrelation in actual interest rates. We note that, in some cases, smoothing is facilitated with only one lag (e.g. Clarida et al. 2000), and in other cases, with more than one lag (e.g. Coibion and Gorodnichenko, 2011).

In our setting, we directly estimate the parameters β_π and β_y in i_t^* after accounting for the dynamics of actual interest rates, inflation, and output. In effect, we augment the above basic specification for the actual interest rate by accounting for lags in the entire system of variables x_t. To understand this, note that the model for actual interest rates, inflation, and the output gap that we estimate can be rewritten in line with equation (3), which accounts for lags between the variables in x_t. If the state of the system follows $S_t = 2$ then the rank of the system is equal to unity, and the resulting multivariate system for x_t can be expressed using equation (10), which we repeat below.

$$\left[i_t - i_t^*\right] \alpha_t = \left[i_t - \beta_{\pi,t}\pi_t - \beta_{y,t} y_t\right] \alpha_t = x_t B_t^*(L) - c_{t+1} + \epsilon_t^*.$$

Note that in the basic equation (16), the ρ parameter reflects convergence to i_t^*. As such, actual interest rates will only instantaneously adjust to target rates if $\rho = 0$. In our setting, both the α_t and B_t coefficients influence the adjustment to i_t^*. Moreover, the adjustment to the target rate can vary over time, hence, the model allows for time-varying smoothing.[9] Figure 1 below shows the adjustment to $i_t^* = 1$ when initial nominal rates are at -1% (essentially interest rates going from 1% below the mean to 1% above the mean) based on the posterior means of the estimated α_t and B_t parameters over all time periods.[10] We discuss the time-variation in the level of smoothing produced by our model in Section 3.4.

3. Results

3.1. Model estimation

The model is estimated using quarterly US output, inflation, and interest rate data from 1954Q3 to 2019Q2. We use the CBO measure of real potential GDP to construct our measure of the output gap (also used in Clarida et al. 2000), the annualized percentage change in CPI as the measure of inflation, and the Federal Funds rate as the nominal interest rate measure. All data were downloaded from the FRED database maintained by the Federal Reserve Bank of St. Louis and demeaned prior to estimation.[11]

The individual steps involved in model estimation are described in Appendix B. Further details regarding the estimation of models with both time-varying rank conditions and time-varying

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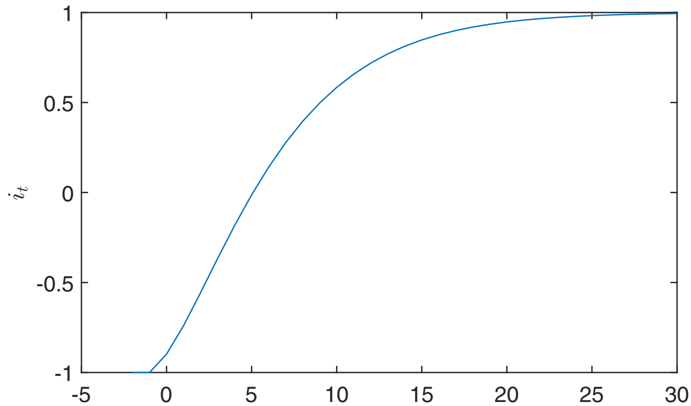

Figure 1. Convergence of actual interest rates i_t to the target rate $i_t^* = 1$ when using the posterior means of the parameter estimates.

parameters are provided by Chua and Tsiaplias (2018). We estimate the model with both the unadjusted Federal Funds rate and the Federal Funds rate adjusted for the Zero Lower Bound (ZLB). With the exception of differences over the period 2009 to 2015, the results are similar using either the adjusted or unadjusted interest rates. In terms of the latter, to deal with the effects of the ZLB on the Federal Funds rate, we adopt the common approach of substituting the "shadow" interest rate into the Federal Funds rate over the period 2009Q1 to 2015Q4 (Krippner, 2015). The objective of the shadow rate measure is to identify what the Federal Funds rate would have been over this period (taking into account the unconventional monetary policy interventions that took place) in the absence of the ZLB. Since the results are similar across the two interest rates, we present only the results associated with the ZLB-adjusted Federal Funds rate.

The parameters, provided in Appendix C, are based on the output from 250,000 draws of the sampler described in Appendix B. We discard the first 50,000 draws. Convergence is fairly quick and we have confirmed convergence to the same posterior density using different initial parameter values.

To determine the model's capacity to reliably fit the data, we focus on the credible intervals produced by the model. Appendix D shows that the credible intervals stemming from the model accurately reflect the time-variation in interest rates, inflation, and output. The results show that the credible intervals are highly efficient, with the model producing an accurate characterization of the time-varying dynamics inherent in the system.

Although our results are based on the preferred CBO-based output gap measure, we have also estimated the model using linearly detrended log real GDP per capita. Our results regarding the level of monetary policy activism remain largely unchanged, with the discussion reserved for Appendix E.

Before examining the time-varying monetary policy weights, we note that, in all cases, the probability of being in the third or fourth regimes (whereby $r = 2$ or $r = 3$ orthogonal, linear contemporaneous relationships are identified in x_t) is close to zero. Consequently, at most, a single contemporaneous, linear relationship can be reliably identified between the variables in x_t. This relationship is expressed as equation (10) and is observationally equivalent to a Taylor-type rule whereby monetary policy is set by targeting a combination of inflation and output.

It is, however, also clear that there are periods where no long-run relationship can be identified in the data. This is apparent from Figure E2 in Appendix E where it is shown that the time-varying probability associated with $r = 1$ (viz. $S_t = 2$) is not always equal to unity. Since the probabilities associated with $r = 2$ or $r = 3$ (viz. $S_t = 3$ or 4 respectively) are close to zero, the time-varying probability of $r = 0$ can be deduced from Figure E2 as $(1 - P(S_t = 2))$. Accordingly, the data are

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not consistent with the standard practice of an a priori imposition of a Taylor-type relationship (in other words, imposing the restriction that the probability of $r = 1$ is always equal to unity). The empirical implications of relaxing this restriction in terms of the estimated monetary policy weights are described in the next section, which compares the monetary policy weights obtained with and without the $r = 1$ restriction.

3.2. The time-varying monetary policy weights and the implications of imposing rank $r = 1$

A key benefit of the analysis in this paper is that we obtain estimates of the monetary policy weights without imposing the restriction that the rank of the system is equal to unity. To examine the ramifications of the relaxation of the rank restriction, we estimate our model both with time-varying ranks and with the restriction that the rank is equal to unity for all time. The latter is consistent with the approach in the existing literature.

Another benefit of our approach is that we are able to monitor the evolution of monetary policy weights over the entire time period, without the constraints stemming from having to a priori identify specific time periods of interest. This is important, given evidence that the approach of dividing sample periods into pre-Volcker and Volcker-Greenspan periods is potentially misleading for examining monetary policy (Kim and Nelson, 2006).

Figure 2 shows significant time-variation in the inflation and output targeting parameters ($\beta_{\pi,t}$ and $\beta_{y,t}$, respectively) over the period 1955 to 2019. Moreover, it is clear that the imposition of the restriction that rank is always equal to unity (i.e. $S_t = 2$ for all t) has significant ramifications for the estimated monetary policy weights.

The results show that, after a relatively long period of stability in the inflation-targeting parameter, a swing in the weight attached to inflation is observed from the late 1970s up until 1985 (Figure 2a). The general rise in the weight attached to inflation from the late 1970s is also observed in the time-varying parameter model presented in Boivin (2006). Consistent with our findings, Boivin (2006) provides evidence that the response of interest rates to inflation increased rapidly from 1979, rising until the mid-1980s, with an estimated weight of approximately 2. The swing, however, appears inordinately strong in the case of $r = 1$. This is primarily due to the $r = 1$ restriction masking the instability in the relationship between interest rates, inflation, and output observed during this period. When the $r = 1$ restriction is permanently imposed, the instability between the key variables is mistakenly absorbed by the inflation and output targeting weights.

A second swing in the inflation-targeting weight appears to commence in 1997, during the Asian financial crisis. This swing is entirely missing from the inflation-targeting weights based on the model with $r = 1$. The swing appears to persist until 2004, declining thereafter. The weights associated with inflation are inordinately large during this period, with monetary policy depending heavily on inflation rather than output. This emphasis on inflation is not observed when restricting the rank to unity, indicating the large impact of instability on the weights associated with inflation during this period. Wieland and Wolters (2013) also present evidence that weights on inflation may be much higher than often estimated, noting that, in a series of forecasting exercises, the optimal inflation response coefficient was almost always close to or at the upper limit of the [0,3] interval that they allowed for. The importance of stability is also highlighted in Bunzel and Enders (2010), who present evidence of non-linearities in the Taylor rule culminating in varying rates of monetary policy aggression. In line with the aforementioned papers, we find that robust estimates with respect to uncertainty about parameters and cointegration can lead to substantially higher weights than typically estimated. Conversely, when we do not account for rank instability, our estimated weight on inflation for the relevant period falls to 2.69 (which is close to the 2.55 estimated by Clarida, et al. (2000) for the period 1982 to 1997 and similar to the 2.19 estimate in Lubik and Schorfheide (2004)).

In general, allowing for rank changes implies that (from the mid-1990s onwards) monetary authorities engaged in far more aggressive monetary policy than that estimated when the rank is

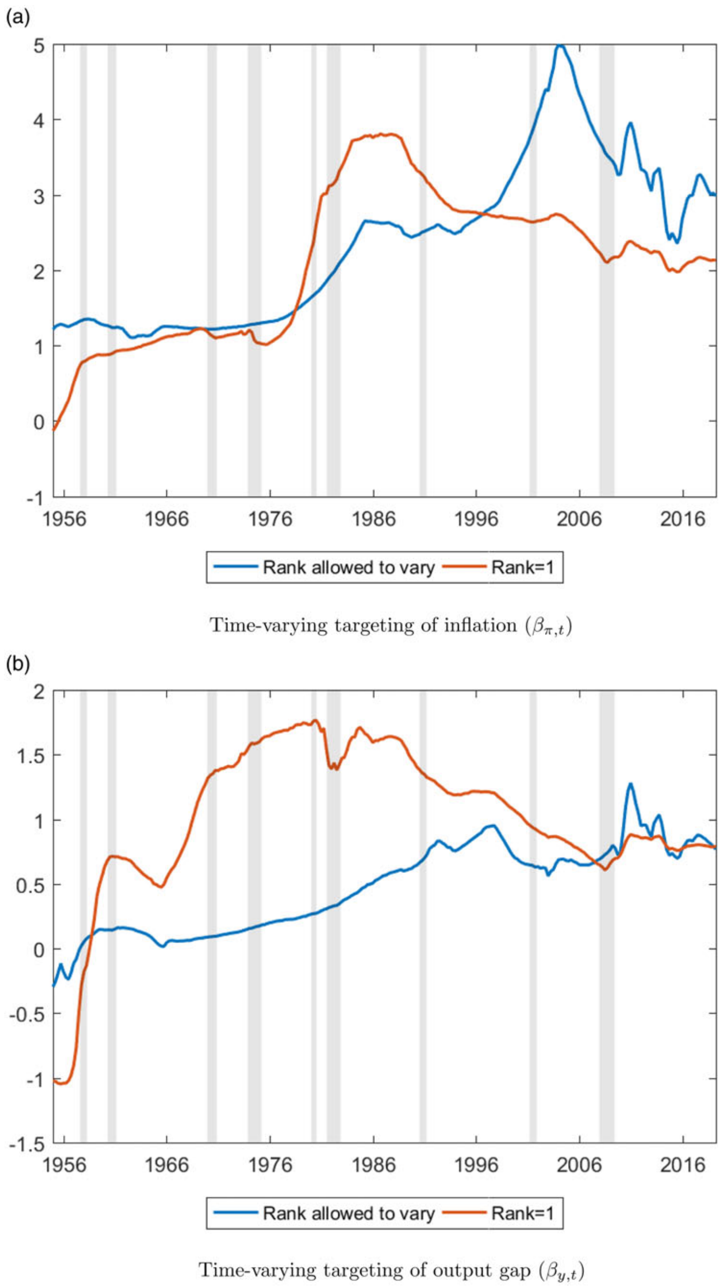

Figure 2. Time-varying targeting of (a) inflation ($\beta_{\pi,t}$) and (b) output gap ($\beta_{y,t}$) for the US effective funds rate from 1955 to 2019, both with and without the restriction that rank = 1. Shaded lines are NBER-dated recessions.

forced to equal unity. The difference in the pre-GFC inflation-targeting weights based on the models with and without rank restrictions is particularly interesting from the perspective of Taylor's (2007, 2012) argument that the absence of sufficient inflation targeting helped contribute to the GFC. The results in the case of $r = 1$ tend to support Taylor's argument, given that the weights attached to inflation are significantly below those estimated in the 1980s. However, the estimates obtained without the $r = 1$ restriction show little support for Taylor's argument, with the estimates indicating that, after accounting for time-varying uncertainty about the presence of a

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common stochastic relationship between the key variables, monetary policy was clearly responsive to inflation over the period 2000 to 2006.

Figure 2b also shows that the imposition of $r = 1$ has a significant impact on the weight attached to the output gap. The gap between the two weights was particularly onerous during the 1970s, with the condition $r = 1$ suggesting that monetary policy attached a larger weight to the output gap during the 1970s. During this period, there was a high probability of a breakdown in the relationship between interest rates, inflation, and output, which likely influences the weight attached to the output gap when the breakdown is not accounted for. This is discussed further in Section 3.7. Figure 4b in Section 3.7, in particular, shows that the time-varying probabilities of $r = 1$ (i.e. $S_t = 2$) are fairly close to zero during the period spanning the late 1960s and the 1970s. This is consistent with evidence in Ahmed et al. (2004) regarding the presence of large exogenous shocks over the period 1960 to 1983 that impacted heavily on US macroeconomic stability. Since rank breakdowns cannot be accommodated in the case where $r = 1$ is imposed prior to estimation, excess weight is attached to the output gap in an attempt to fit the data.

To better glean the impact of rank restrictions on the relative weights attached to inflation and the output gap, Figure F1 in Appendix F presents the time-varying difference between the weights on inflation and the weights on the output gap when estimation allows for the rank to vary and when the rank is permanently set to unity. The results suggest that the $r = 1$ restriction tends to generate a relative weight in favor of the output gap during the 1970s, hence suggesting a (relative) under-targeting of inflation which disappears in the early 1980s. This property observed when setting $r = 1$ is similar to that in Clarida et al. (2000), who found that variation in the inflation-targeting coefficient pre- and post- Volcker was very large, whereas variation in the output targeting coefficient was much smaller.[12] In the absence of rank restrictions, the results continue to show a relative increase in the weight attached to inflation during the Volcker regime. However, the increase is materially smaller than that estimated when we impose $r = 1$ (viz. imposing $r = 1$ tends to under-emphasize the adverse impact of stability-related shocks during the 1970s, thereby over-emphasizing the purported absence of monetary policy). Coupled with the evidence of severe rank instability during the 1970s, the results tend to highlight the particular importance of adverse shocks in explaining economic conditions during the Great Inflation of the 1970s (Sims and Zha, 2006; Ahmed et al. 2004).

The results also indicate that the failure to allow for time-variation in the rank of the system governing the variables in x_t has substantive ramifications for understanding the evolution of monetary policy weights. The a priori (largely implicit and little discussed) imposition of $r = 1$ leads to conclusions that conflate the monetary policy weights with the impact of shocks that influence the stability of the relationship between interest rates, inflation, and output. After allowing for time-variation in the rank of the system governing the key variables of interest, there is relatively little evidence in favor of the argument that the relative under-targeting of inflation contributed materially to the Great Inflation (which, as discussed further in Section 3.7, is largely characterized by a persistent breakdown in the relationship between interest rates, inflation, and output) or the argument that monetary policy inadequately targeted inflation prior to the Global Financial Crisis.

3.3. Constructing measures of the time-varying probability of activism and determinacy

To study the time-varying probability of active or passive monetary policy, we focus on the parameter estimates for the inflation-targeting coefficient, with monetary policy being passive if $\beta_{\pi,t} < 1$.[13] The premise behind this notion of passive monetary policy is Taylor's (1993) principle that if nominal rates rise by less than one-for-one with inflation (i.e. monetary policy targets inflation passively such that $\beta_{\pi,t} < 1$) then self-fulfilling cycles are possible.

A key difference between our approach and the approach used historically is that we do not impose the restriction that a Taylor-type rule is permanently present in the data (viz. that $P(S_t = 2) = 1$) for all t). The empirical measurement of monetary policy activism therefore accounts for

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both the pass-through from inflation to interest rates ($\beta_{\pi,t}$) and the evidence in favor of a single common relationship between interest rates, inflation, and output ($S_t = 2$).

The primary issue with relying only on $P(\beta_{\pi,t}) > 1$ to assess the probability of activism is that it conflates the desired outcome with the realized outcome, potentially providing false comfort about the probability of monetary policy activism. For example, relying only on $P(\beta_{\pi,t}) > 1$ allows for statistically anomalous situations such as a very high probability of activist monetary policy notwithstanding $P(S_t = 2)$ being small or even close to zero.

To account for the possibility that the rank of the system is not equal to unity, and thereby correct for the above issue, our probability of active monetary policy at time t is

$$P(activism_t) = P(\beta_{\pi,t} > 1)P(S_t = 2) \tag{17}$$

whereby active monetary policy requires conditions on both the pass-through of inflation and the presence of a relationship between the three key variables.[14]

Similarly, we can identify the time-varying effective probability of determinate monetary policy taking into account the possibility that the rank of the system governing interest rates, inflation, and the output gap is not always equal to unity. This results in an effective probability of determinate monetary policy at time t of

$$P(determinacy_t) = P(BK_t)P(S_t = 2) \tag{18}$$

where $P(BK_t)$ is the probability of satisfying the Blanchard–Kahn (hereafter denoted BK) conditions at time t. The failure to satisfy the BK conditions implies that self-fulfilling cycles (viz. sunspot-type solutions) are possible.

We, therefore, consider the probability of determinacy in light of two stability conditions: (i) the stability associated with the satisfaction of the Blanchard–Kahn conditions (which are contingent on parameters that are only identified in the presence of a common relationship between the variables in x_t); and (ii) the stability associated with the probability at time t of observing a common stochastic relationship between interest rates, inflation, and the output gap. In the absence of the latter, the parameters in the former are not identified.

We note that the first condition in equation (18) (associated with $P(BK_t)$) is standard in existing research (Clarida, et al. 2000; Lubik and Schorfheide, 2004; Cogley and Sargent, 2005). The second condition is, however, implicitly omitted by way of the a priori setting of $P(S_t = 2) = 1$. This is not without empirical consequence, however, as there are many periods where $P(S_t = 2)$ is clearly not equal to unity (these periods are discussed further in Section 3.7).

3.4. Monetary policy activism versus monetary policy determinacy

To link the results to monetary policy determinacy, we first consider a business cycle model with sticky prices in the vein of that used in Clarida et al. (2000). Implicit in this model is the absence of trend inflation. After discussing the results for the no-trend inflation scenario, in Section 3.8, we use the model proposed by Coibion and Gorodnichenko (2011) to evaluate the impact of trend inflation on our estimated time-varying probabilities of monetary policy determinacy.

The log-linearized model around a zero-inflation steady state yields the following equilibrium relationships

$$\pi_t = \delta E_t(\pi_{t+1}) + \lambda y_t \tag{19}$$

$$y_t = \beta E_t(y_{t+1}) - \frac{1}{\sigma}(i_t - E_t(\pi_{t+1})) \tag{20}$$

$$i_t = (1-\rho)i_t^* + \rho i_{t-1} + e_t \tag{21}$$

where i_t^* follows a Taylor-rule type structure as in equation (15).

We use the draws of $\beta_{\pi,t}$ and $\beta_{y,t}$ from our model to determine the proportion of draws for which the BK conditions are satisfied, thereby yielding an estimate of the probability of determinacy in each time period. The failure to satisfy the BK conditions allows for self-fulfilling expectations and implies monetary policy indeterminacy. In effect, we follow the same approach as in Clarida et al. (2000) and Coibion and Gorodnichenko (2011), feeding the empirical Taylor rule estimates from our model into a theoretical model. As noted in the latter paper, a benefit of this approach is that it accounts for possible misspecification of the theoretical model (which is particularly important in a long-time setting such as the one in our paper).

To evaluate whether the BK conditions are satisfied, we require estimates of the smoothing parameter ρ. In this respect, we set ρ such that the number of periods it takes to converge to a target i_t^* is identical to that implied by the values of a_t and B_t, hence using either ρ or a_t, B_t yields convergence to the target in the same number of periods.[15] The resulting ρ reflects the level of interest rate smoothness estimated by our model. We note that the adoption of the posterior means of a_t and B_t over all time periods yields a level of smoothing that is equivalent to $\rho \approx 0.87$. Appendix G shows the implied smoothness parameter ρ (or, more precisely, ρ_t) produced by our model in each time period. Consistent with the implied smoothness in Coibion and Gorodnichenko (2011) (which the authors construct as the sum of AR(1) and AR(2) coefficients), we see a decline in smoothness during the 1970s and an increase in the level of inertia during the 1980s. Thereafter, inertia in actual interest rates has remained elevated. A key difference between our estimates and the time-varying estimates in Coibion and Gorodnichenko (2011) is that the smoothness estimated during the 1970s in the latter is approximately 0.6, whereas our smoothness is approximately 0.75. In contrast, during the period 1980 to 2002, Coibion and Gorodnichenko estimate an average smoothness of about 0.85 (which is similar to our estimate of about 0.9). Overall, our model indicates a high level of inertia in actual interest rates, thereby supporting the notion that the Federal Reserve engages in smoothing when adjusting interest rates. Finally, we follow Clarida et al. (2000) in setting the discount factor δ to 0.99, the output elasticity of inflation λ to 0.3, and the coefficient of relative risk aversion σ to 1.

Figure H1 in Appendix H compares the time-varying probability of determinate monetary policy with the time-varying probability of monetary policy activism (viz. $P(\beta_{\pi,t} > 1)$). Values of $\beta_{\pi,t}$ that are below unity can still yield determinate monetary policy, but only if $\beta_{\pi,t}$ is only slightly below unity (e.g. 0.99). The ρ parameter influences the upper bound of the values that $\beta_{\pi,t}$ can take (e.g. very large values of $\beta_{\pi,t}$ can yield indeterminate monetary policy). However, the upper bound is higher than almost every value of $\beta_{\pi,t}$ that we draw, hence, the results are effectively invariant to the value of ρ. Accordingly, although the probabilities of activism and determinacy differ slightly, the difference is negligible. We note that, since $P(BK_t)$ and $P(\beta_{\pi,t} > 1)$ are almost identical, it follows that, for the model specified in equations (19)–(21), $P(determinacy_t)$ in equation (18) is essentially identical to $P(activism_t)$ in equation (17). This no longer holds, however, when we allow for trend inflation in Section 3.8.

3.5. Time-variation in the determinacy probabilities

A key feature of our approach is that we decompose the probability of determinacy into two stability considerations: (i) the stability based on the satisfaction of the Blanchard–Kahn conditions; and (ii) the stability underpinning the notion that interest rates, inflation, and the output gap share a common stochastic relationship (which is implicitly set to unity in the literature on monetary policy determinacy and activism). By setting the latter to unity, it is therefore not possible to determine the extent to which the probability of determinacy is influenced by both stability considerations (essentially, as we see in Section 3.2, contaminating the estimation of the parameters that are needed to determine whether the Blanchard–Kahn conditions are satisfied). Examining time-variation in $P(BK_t)$ and $P(S_t = 2)$ (or, in the case of activism, $P(\beta_{\pi,t} > 1)$ and $P(S_t = 2)$) is therefore of primary importance.

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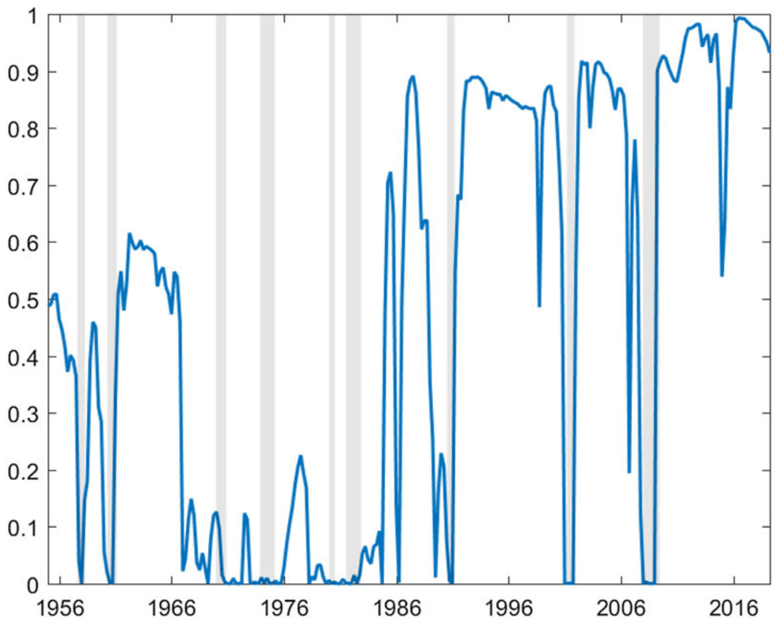

Figure 3. Time-varying probability of monetary policy determinacy. The probability is based on equation (18). Shaded lines are NBER-dated recessions.

Before considering the individual factors $P(BK_t)$ and $P(S_t = 2)$, we examine the overall probability $P(determinacy_t)$ (which is the product of $P(BK_t)$ and $P(S_t = 2)$). The results in Figure 3 show that, from 1955 to 2019, the probability of determinate monetary policy in a given quarter has varied considerably, ranging from close to zero to close to unity. During the extended period spanning 1967 to the early 1980s, the probability of determinate monetary policy tended to be between 0% and 20%. From the mid-80s, we observe a marked shift in the probability of determinate monetary policy, with the probability of determinate monetary policy hovering at about 70 percent. This value is materially influenced by the presence of many periods of instability, which lower the average probability of determinacy and impede the capacity to target inflation.

Taken at face value, our results regarding the determinacy of monetary policy during the 1970s are broadly consistent with typical findings that identify an increase in the probability of determinacy in the 1980s (Clarida, et al. 2000; Lubik and Schorfheide, 2004; Cogley and Sargent, 2005; Coibion and Gorodnichenko, 2011; Hirose, et al. 2020). However, as noted above, the aforementioned papers do not allow for instability in the presence of a common relationship between the variables in x_t. In Section 3.7, we show that such a distinction is critical for understanding the factors driving the determinacy of monetary policy (and monetary policy activism). In particular, we show that estimates of monetary policy indeterminacy in the 1970s are largely attributable to instability in the rank of the relationship between interest rates, inflation, and output rather than due to either $P(\beta_{\pi,t} > 1)$ or the probability of satisfying the BK conditions.

In terms of the determinacy of monetary policy in the period preceding the Great Recession, our results tend to differ from Doko-Tchatoka et al. (2017), who present evidence of indeterminate monetary policy in the period between the 2001 slump and the onset of the Great Recession (specifically 2002Q1 to 2007Q3). In this respect, Doko-Tchatoka et al. (2017) assume zero-trend inflation, so we compare their results with our model when we also assume zero-trend inflation (rather than the results we obtain when we allow for trend inflation in Section 3.8). On average, we estimate that the average probability of determinate monetary policy during this period was approximately 80%, thereby providing little evidence in support of indeterminate monetary

Table 1. Average probability of determinate monetary policy

	LS, 2004	TVP-R
Pre-Volcker	0.00	0.20
Post-1982 (to 1997Q4)	0.98	0.56
Volcker-Greenspan (to 1997Q4)	0.38 to 0.70	0.46
All data to 2019Q2		0.48
All data to 2019Q2 (excl. recessions)		0.55
Post-1990		0.75
Post-1990 (excl. recessions)		0.83

Notes: LS, 2004 refers to the estimates in Lubik and Schorfheide (2004). TVP-R refers to the estimates based on the model in this paper. The recession indicator is set to unity if any month in the quarter was deemed a recession according to the NBER. Pre-Volcker is 1960Q1–1979Q4. Volcker-Greenspan is from 1979Q3:1997Q4.

policy. However, we also find that the determinacy probabilities decline markedly toward the end of 2006, averaging 57% from 2006Q4 to 2007Q3.[16] This raises the possibility of passive monetary policy in the period shortly before the onset of the Great Recession but rejects the more important notion of a sustained period of indeterminate monetary policy between 2001 and 2007.

3.6. Monetary policy determinacy during key historical periods

It is useful to compare our estimates with those obtained by others during key periods of interest. Clarida, et al. (2000) and Lubik and Schorfheide (2004), for example, differentiate between the pre- and post-Volcker periods in examining monetary policy determinacy. We focus on the probabilities in Lubik and Shorfheide, which broadly reflect popular findings on determinacy during the 1970s and 1980s (e.g. Clarida, et al. 2000; Cogley and Sargent, 2005; Coibion and Gorodnichenko, 2011).

The posterior probabilities estimated by Lubik and Shorfheide in Table 1 reveal striking differences in the probability of determinate monetary policy in the 1970s and 1980s. These can be compared to the probabilities in our model, which we denote in Table 1 as the 'TVP-R' model. The three sub-periods presented in the table are based on the periods chosen by Clarida, Gali, and Gertler and Lubik and Shorfheide: a pre-Volcker period from 1960Q1 to 1979Q2; a Volcker-Greenspan period from 1979Q3 to 1997Q4, and a post-1982 sample from 1982Q4 to 1997Q4. We also consider the overall probabilities across the entire sample and probabilities post-1990 both with and without recessions.

The pre-Volcker period in Lubik and Schorfheide concentrates (almost) all of its probability mass in the indeterminate monetary policy region. Most other papers find a similar result. The evidence regarding the probability of monetary policy determinism during the Volcker-Greenspan sample period is, however, mixed. Depending on the choice of prior, the probability estimates of determinate monetary policy in Lubik and Schorfheide range from 0.38 to 0.70.[17] The authors argue that this variation is possibly influenced by the Volcker disinflation period, which may be better characterized by non-borrowed reserve targeting than by an interest rate rule. In so doing, they contend that if this disinflation period is excluded, the posterior probability of determinate monetary policy is close to unity for the post-1982 sample. Accordingly, the probabilities of determinate monetary policy pre- and post-Volcker are essentially binary (going from zero to unity).

To obtain period-specific estimates, we take the sample mean of our quarterly probabilities of monetary policy determinacy during the periods of interest. For the post-1982 period, both Lubik and Schorfheide and our model suggest a higher probability of determinate monetary policy, although our probability is well below unity. In general, our estimates provide little support for the general finding of binary probabilities (e.g. from indeterminacy with probability 1 to determinacy with probability 1). Instead, there is almost always a substantive level of uncertainty about the determinacy of monetary policy. Relatedly, the incorporation of time-varying uncertainty about the monetary policy parameters and about the stability of the relationship between interest rates, inflation, and output tends to reduce the average probability of determinacy, producing probabilities that are materially lower than unity; over the period post-1982 to 1997 the average probability of determinacy is 56%, with the average probability of determinacy in the post-1990 period being 75%.

3.7. Time-variation in the factors underpinning monetary policy activism and determinacy

This section augments the preceding analysis by estimating, for the first time, how monetary policy determinacy can be decomposed into: (i) the component reflecting the distribution of the monetary policy weights and the satisfaction of the BK conditions; and (ii) the component reflecting instability in the presence of a common stochastic relationship between interest, rates, inflation, and output. In so doing, we show that much of the estimated indeterminacy in monetary policy in the 1970s is attributable to instability (whereas seminal papers generally assume that the estimated passivity is a reflection of the distribution of the monetary policy weight on inflation or the presence of trend inflation).

A key issue in previous research is that probabilities regarding monetary policy determinacy and activism are conditional on the assumed continuous presence of a fixed monetary policy rule. For example, Clarida, et al. (2000), Lubik and Schorfheide (2004), and Cogley and Sargent (2005) assume that a Taylor-type rule must hold in all time periods. This is equivalent to assuming that $P(S_t = 2) = 1$ for all t. By relaxing this restriction, a benefit of our approach is that we are able to identify time-variation in both the probability of satisfying the BK conditions (or, in the case of activism, the probability of $\beta_{\pi,t} > 1$) and the probability that $S_t = 2$.

Figure 4 shows significant time-variation in both the probability of satisfying the BK conditions and in the probability of a single common relationship between interest rates, inflation, and output (viz. $P(S_t = 2)$). Pursuant to Figure 4a, the probability of satisfying the BK conditions is approximately 60 percent over the period 1955 to the late 1970s, rising to 71 percent by the end of 1980 and above 80 percent from 1983. Figure 4b highlights the importance of accounting for time-variation in the rank of the system of variables when attempting to determine the probability of determinate monetary policy. The figure clearly shows periods of a breakdown in the relationship between interest rates, inflation, and output, particularly during the period 1967 to the early 1980s. The figure also shows the temporary breakdown of the relationship between output, interest rates, and inflation during recessions and downturns. This can be contrasted with the persistent breakdown from the late 1960s to the early 1980s.

We note that evidence presented in Figure 4b indicating a high probability of $r = 0$ from the late 60s through the 1970s and the early 1980s is also consistent with Kahn et al. (2001) and Ahmed et al. (2004). The two aforementioned papers provide evidence of instability over the course of the 1970s and early 1980s, with a shift observed around 1984 due to factors such as technology and the fortuitous absence of large economic shocks.

In general, we observe ongoing improvement in the probability of both active monetary policy and the probability of satisfying the BK conditions over the sample period. The average probability of the latter rises from 20% pre-Volcker to 56% from 1983 to 1997. Post-1990, the average probability of satisfying the BK conditions rises to 75%. The estimates in Figure 4 show that this pattern of rising activism over the past three decades is due to both changes in monetary policy, resulting in a rising probability in favor of satisfying the BK conditions (and yielding $\beta_{\pi,t} > 1$), and the absence of shocks such as those observed in the 1970s (that resulted in close to zero values for $P(S_t = 2)$ for much of the 1970s).

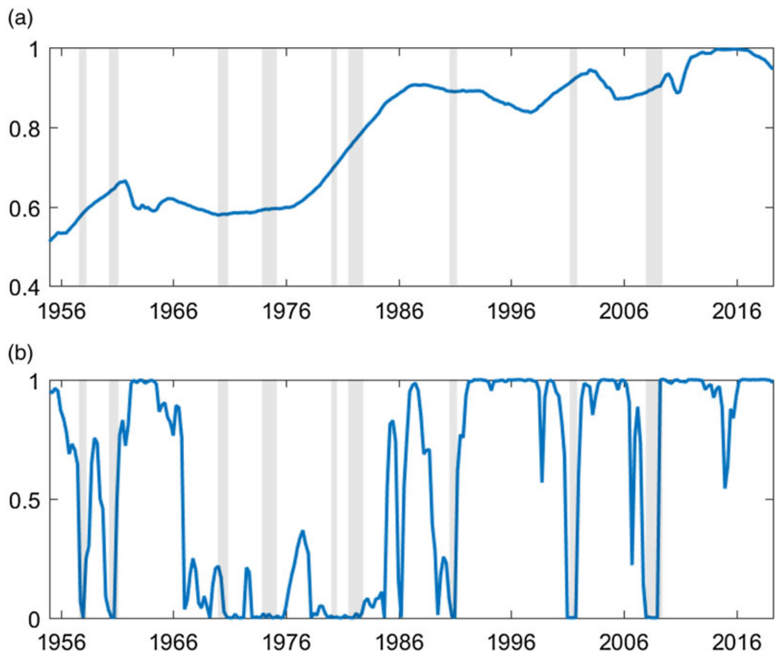

Figure 4. Time-varying probability of monetary policy determinacy factors. Figure (a) is the time-varying probability of satisfying the Blanchard–Kahn conditions. Figure (b) is the time-varying probability of $S_t = 2$ (i.e. rank = 1), whereby a single linear relationship between observed interest rates, inflation, and output is identified at time t. Shaded lines are NBER-dated recessions.

Importantly, our estimates of generally indeterminate (and passive) monetary policy during the 1970s are primarily attributable to small values of the probability of $S_t = 2$ (viz. due to instability) rather than to the probabilities associated with the satisfaction of the BK condition or the extent to which $P(\beta_{\pi,t} > 1)$. In so doing, the results reject the widely-held position (e.g. Clarida et al 2000; Lubik and Schorfheide, 2004) that a failure to respond sufficiently to either inflation or the output gap during the 1970s materially contributed to the inflation observed during this period. Given the breakdown in the relationship between interest rates, inflation, and output, even if the probability of satisfying the BK conditions (or observing $P(\beta_{\pi,t} > 1)$) was significantly greater than that observed during the 1970s, the effective probability of determinate monetary policy would continue to have been relatively small. The results highlight the critical importance of adverse shocks (viz. fortuitous circumstances) for the effectiveness of monetary policy to reign in inflation, which can be mistakenly attributed to overly passive monetary policy.

3.8. Determinacy and trend inflation

Coibion and Gorodnichenko (2011) find that, with non-zero-trend inflation, the probability of determinacy cannot be approximated using the rule $P(\beta_{\pi,t} > 1)$. To examine the ramifications of trend inflation for the determinacy of monetary policy, we, therefore, consider the New Keynesian model with Calvo pricing and non-zero-trend inflation used in Coibion and Gorodnichenko (2011). To preserve space, we do not reproduce the model and refer readers to the aforementioned paper. In their paper, the authors estimate time-varying parameters of monetary policy weights attached to inflation, the output gap, and output growth and consider the determinacy of monetary policy under alternative levels of trend inflation by feeding draws of their parameter estimates into their structural model and determining the average probability of monetary policy determinacy at each time period.

Prior to undertaking a similar exercise, we note that a particular property of the model in Coibion and Gorodnichenko is that a stronger weight on the output gap can severely reduce the probability of determinacy. For example, consider the distribution of our β_π draws at a random date 1984Q1 and set the output gap and output growth values to $\beta_y = 0.8$ and $\beta_{gy} = 1$, respectively. Our posterior mean of β_π in this period is 2.3. We assume trend inflation of 3% and, for the other parameters in the model, adopt the same values as in Coibion and Gorodnichenko. The probability of determinacy is approximately 30% (using either a contemporaneous or forward-looking monetary policy rule). However, if β_y is then reduced to 0, the probability of determinate monetary policy rises to 82%. The same is observed even if we set the weight on output growth to 0, hence, the inclusion or omission of output growth does not negate this property.

A practical consequence of the above is that, under the trend inflation model, the monetary authority can generate extreme jumps in the probability of determinacy by simply ignoring the output gap. These jumps are not readily evident in Coibion and Gorodnichenko because their estimates of β_y are fairly constant (at between 0.4 and 0.5) over the entire period from 1969 to 2002. Because our model produces large changes in the targeting of the output gap (from the 1990s onwards, our estimates of the weight on the output gap are double those of Coibion and Gorodnichenko), the implied probability of determinate monetary policy declines simply because of the stronger targeting of the output gap. As noted above, these results are not attributable to the inclusion or omission of output growth, and we obtain similar determinacy probabilities when we set the weight on output growth to 0, 1, or 2 (hence essentially spanning the different weights on β_{gy} used in Coibion and Gorodnichenko).

To facilitate a comparison with the relatively constant β_y in Coibion and Gorodnichenko, we determine the probability of determinacy under 3% trend inflation with a contemporaneous monetary policy rule using our time-varying draws of β_π, but replacing our time-varying β_y with its unconditional mean of approximately 0.45. We then estimate the probability of determinacy using both $\rho = 0.87$ (which is the average ρ implied by our model) and the time-varying values of ρ in Appendix G, obtaining similar results. Moreover, we estimate probabilities of determinacy using output growth weights of $\beta_{gy} = 0$ and $\beta_{gy} = 1$ with similar results. Hence, we limit presentation of our results to those based on $\rho = 0.87$ and $\beta_{gy} = 1$. The remaining parameter values are those used by Coibion and Gorodnichenko in their 2011 analysis.[18]

Figure 5a shows the estimated probabilities of satisfying the BK conditions using both zero-trend inflation (which are identical to those in Figure 4a) and those based on 3% trend inflation. Figure 5b shows the probabilities of determinacy after accounting for the time-varying probability of identifying a common stochastic relationship between interest rates, inflation, and output across the zero-trend inflation and non-zero-trend inflation models. Using the time-varying posterior distributions of our inflation weights β_π, we find that shifting from zero-trend inflation to 3% trend inflation tends to reduce the probability of satisfying the BK conditions by an average of 15 percentage points. In some periods, the difference in the probability of satisfying the BK conditions is close to zero. However, the difference exceeds 30 percentage points in certain periods (namely in 1992 and 2014). When we account for the possibility that $P(S_t = 2)$ can be less than unity (Figure 5b), we estimate that the average probability of determinacy when allowing for typical trend inflation falls by approximately 10 percentage points (over the period 1955 to 2019).

To better glean the average probabilities under zero and non-zero-trend inflation, in Table 2, we reproduce the determinacy probabilities that we constructed in Table 1 focusing on the probability of determinacy across zero and 3% trend inflation. The possibility of non-zero-trend inflation further reduces the probability of determinate monetary policy. However, irrespective of zero or non-zero-trend inflation, we continue to observe that the probability of determinacy is materially depleted when we relax the assumption that the probability of a permanent stochastic relationship (viz. a relationship that can be identified at each time period) between interest rates, inflation, and output is equal to unity. In particular, irrespective of the presence of a trend in inflation, our

Table 2. Probability of determinate monetary policy with zero and non-zero-trend inflation

	Zero-trend	3% Trend
Pre-Volcker	0.20	0.16
Post-1982 (to 1997Q4)	0.56	0.41
Volcker-Greenspan (to 1997Q4)	0.46	0.34
All data to 2019Q2	0.48	0.38
All data to 2019Q2 (excl. recessions)	0.55	0.44
Post-1990	0.75	0.59
Post-1990 (excl. recessions)	0.83	0.65

Notes: Pre-Volcker is 1960Q1–1979Q4. Volcker-Greenspan is from 1979Q3:1997Q4.

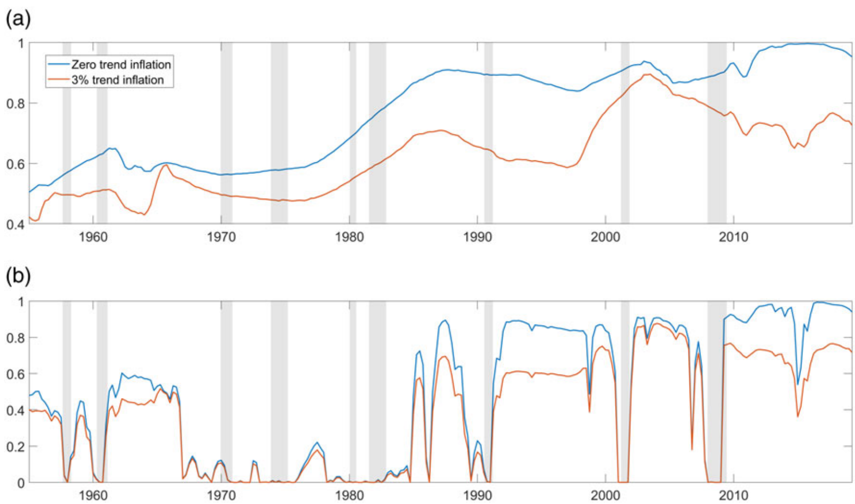

Figure 5. Comparison of determinacy under zero and non-zero-trend inflation. Figure (a) is the time-varying probability of satisfying the Blanchard–Kahn conditions with zero-trend and 3% trend inflation. Figure (b) is the time-varying probability of $P(determinacy_t)$ taking into account uncertainty about whether a common relationship between interest rates, inflation, and output is identified at time t. Shaded lines are NBER-dated recessions.

results continue to reject the notion of binary probabilities regarding determinacy and indeterminacy. Hence, we find a materially greater level of uncertainty about the determinacy of monetary policy than that estimated in seminal papers. For the post-1990 period, the average probability of determinacy is 75% when assuming zero-trend inflation and 59% when assuming 3% trend inflation. Although the average probabilities continue to favor determinacy, they are far below unity and indicate substantially greater average risks of indeterminate monetary policy than typically found in the literature.

4. Concluding remarks

We show that prevailing restrictions in existing research, which imply a permanent relationship (with probability 1) between interest rates, inflation, and output, result in misleading monetary policy weights. In particular, the evidence strongly contradicts the notion of a permanent stochastic relationship between interest rates, inflation, and output. The easing of the aforementioned restriction has significant ramifications for key issues in monetary policy such as the adequacy of inflation targeting and the determinacy of monetary policy.

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New estimates of the probability of monetary policy determinacy are constructed that enable us to decompose the probability of determinacy into the components that are related to: (i) the satisfaction of the Blanchard–Kahn conditions; and (ii) the capacity to identify a common relationship between interest rates, inflation, and the output gap (hence the capacity to identify the parameters that are used to determine satisfaction of the Blanchard–Kahn conditions). The estimates indicate that most of the estimated indeterminacy of monetary policy in the 1970s is attributable to instability (in contrast to the typical argument of inadequate inflation or output targeting). Accordingly, the extent to which more aggressive inflation targeting (or, indeed, stronger targeting of the output gap) would have reigned in inflation during the 1970s is materially overstated.

Overall, our estimates provide little support for the general finding of binary (or close to binary) probabilities of determinacy before and after the early 1980s (e.g. from indeterminacy with probability 1 to determinacy with probability 1). Instead, there is persistently greater uncertainty about whether monetary policy is determinate or indeterminate than generally estimated or assumed in the literature. Moreover, after relaxing the typical restriction of a permanent common relationship between interest rates, inflation, and output (irrespective of whether that relationship is underpinned by constant or time-varying parameters), we estimate materially lower average probabilities of monetary policy determinacy than those in key papers. The lower probabilities are observed under both the zero-trend and non-zero-trend inflation environments.

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Supplementary material. To view supplementary material for this article, please visit https://doi.org/10.1017/S136510052400004X

Notes

1 For simplicity, we assume that all variables are in deviation from their mean.
2 Moreover, a limitation of the research relying on Markovian regime-switching is that it is usually limited to a small number of regimes (for example, dovish and hawkish regimes); as such, a dovish regime in 1955 is no different from a dovish regime in 2005.
3 We do not impose the order of integration of x_t prior to estimation. Instead x_t can be $I(0)$ or $I(1)$ depending on the current behavior of the data. This allows for situations where there is a breakdown in the cointegrating relationship between the variables. Section 2.1 discusses the rank conditions for Θ_t, hence, the order of integration of x_t.
4 Alternatives to time-invariant transition processes are considered by Kim and Nelson (1998) and Kaufmann (2015).
5 Technically, r should have a time subscript but this is omitted for expositional convenience.
6 Chan et al. (2018) also discuss the issue of invariance in ordering for factor models.
7 This expansion is beneficial as it avoids the need to consider a time-varying dimensionality for the parameter space, which would render the computation extremely cumbersome due to the need to employ reversible jump MCMC techniques.
8 This form is broadly in line with that of Taylor (1993), Clarida, et al. (2000), Lubik and Schorfheide (2004), and Boivin and Giannoni (2006). Other specifications have also been adopted, such as the inclusion of output growth (e.g. Coibion and Gorodnichenko, 2011).
9 As a simple example to compare interest rate smoothing in equation (16) with the smoothing in our model, consider $B_t = 0$ and $\alpha_t \in (-1, 0)$. The proposed model will then produce interest rate smoothing equivalent to $\rho = 1 + \alpha_t$.
10 The average level of smoothing implied by the posterior means ($\alpha_t = -0.05$ and a weight of 0.53 on Δi_{t-1}) is equivalent to $\rho = 0.87$, hence similar to the level of smoothing in Coibion and Gorodnichenko (2011).
11 We have explicitly selected the data in line with the variables used in seminal papers such as Clarida, et al. (2000) and Lubik and Schorfheide (2004). However, alternative formulations of Taylor-type rules also exist (e.g. forecast-based interest rate rules). Readers are referred to Wieland and Wolters (2013) for further information.
12 Clarida et al. (2000) note that the estimated output targeting coefficient had a large standard error and (in contrast to the pre-Volcker era) was only marginally significant. Using post-1982 data, they state that "we cannot reject the hypothesis that the Fed has effectively pursued a pure inflation targeting policy."

13 An alternative condition for active policy requires that $\beta_{\pi,t} + \frac{1-\beta}{\lambda}\beta_{y,t} > 1$, where β is the discount rate and λ is the slope of the Phillips curve (Woodford, 2003; Mavroeidis, 2010). The two sets of probabilities, however, are virtually indistinguishable in our model.
14 The time-varying probabilities regarding $P(activism_t)$ can be obtained for each quarter thereby providing unique insights into the periods of passive monetary policy. These probabilities are obtained "exactly" using the draws from the posterior distribution of $\beta_{\pi,t}$ and S_t and are available for every time period.
15 We determine convergence by reference to whether $|i_{t+k} - i^*_t| < 0.01$ is satisfied and set the initial i_t to -1 and i^*_t to 1.
16 Nevertheless, we do not observe a probability of determinacy near the 10% estimated by Doko-Tchatoka et al. (2017) when the latter relies on CPI inflation.
17 This variation also highlights the risks associated with using estimates based on pre-selected periods. For example, the estimates in Lubik and Schorfheide do not provide a clear depiction regarding whether monetary policy was determinate from 1979 onward or only post-1982. If the start date of the period-specific analysis is amended from 1982Q4 to 1979Q3, the model in Lubik and Schorfheide produces extremely different probabilities regarding determinate monetary policy (which go from 0.98 to somewhere between 0.38 and 0.70 and hence favor both determinacy and indeterminacy depending on the chosen probability).
18 Hence, a Frisch labor supply elasticity of 1, a discounting parameter β of 0.99, steady-state growth of 1.5% per annum, an elasticity of substitution equal to 10, and a degree of price stickiness equal to 0.55.

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