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Modest policy interventions ☆

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Abstract

We present a theoretical and empirical framework for computing and evaluating linear projections conditional on hypothetical paths of monetary policy. A *modest policy intervention* does not significantly shift agents' beliefs about policy regime and does not induce the changes in behavior that Lucas (Carnegie–Rochester Conference Series on Public Policy, Vol. 1, Amsterdam, North-Holland, 1976, pp. 104–130) emphasizes. Applied to an econometric model of U.S. monetary policy, we find that a rich class of interventions routinely considered by the Federal Reserve is modest and their impacts can be reliably forecasted by an identified linear model. Modest interventions can shift projected paths and probability distributions of macro variables in economically meaningful ways.

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1. Introduction

This paper presents a framework, which respects the Lucas (1976) critique, to compute and evaluate linear projections of macro variables conditional on hypothetical paths of monetary policy. The framework includes a theoretical model that reports when linear projections are reliable even though policy switches from one regime to another. Insights from the theory are applied to an empirical model of U.S. monetary policy behavior to probe the range of policy interventions that do not significantly change private agents' beliefs about the prevailing policy regime.

We assume that in the true economy policy regime evolves randomly according to a Markov chain where regime cannot be observed by private agents. Endowed with knowledge of the underlying policy process, agents act rationally as Bayesian updaters to infer policy regime from realizations of past policy variables. The true economy generates two objects of interest: (1) nonlinear dynamics with agents updating their beliefs about regime; (2) linear dynamics conditional on a given regime.[1]

A policy advisor, who seeks to inform policymakers of the likely effects of alternative policy actions, performs positive policy evaluation. We assume that due to technological limitations the policy advisor cannot model the true nonlinear economy, so object (1) is unattainable. Instead, we arm the advisor with a (misspecified) linear model, object (2), and consider only exogenous intra-regime changes. We also develop a procedure to check if the alternative policies are perceived by private agents as intra-regime changes. To do this we carefully explore conditions under which linear projections conditional on hypothetical policies will closely approximate the truth.

To estimate a linear model, the advisor selects a sample period when policy has operated in a single regime. Conditioning on that regime, the advisor reports projections conditional on hypothetical intra-regime policy changes. The advisor also provides another useful piece of information: the likelihood that the policy changes trigger shifts in agents' beliefs in the prevailing regime. If the likelihood is high, predictions of policies may be unreliable; otherwise, the projections are likely to be good approximations to the truth.

We call a linear projection under the prevailing regime the *direct effects* and the impacts induced by a change in agents' beliefs about policy regime the *expectations–formation* effects. Direct effects are the usual policy impacts that arise when regime is held fixed. Direct effects do not include any impacts due to changes in agents' beliefs about regime and, therefore, do not shift decision rules. Expectations–formation effects arise when the policy intervention induces agents to alter their beliefs about the prevailing regime. These inter-regime changes shift both the public's expectations–formation rules and their decision rules in the systematic way that Lucas encouraged the profession to model.

[1] One could also obtain large-sample linear dynamics that average across regimes. But this holds little interest because it embodies a linear combination of private decision rules, a combination that is systematically wrong.

Our theoretical framework quantifies the two kinds of effects from policy changes that vary in magnitude and dynamic pattern. We develop a statistic that indicates whether direct effects are improbably large relative to their normal historical variation. An intra-regime policy change is *modest* when this statistic is close to its mean. We show that

- contrary to the priors of many economists, modest policy interventions may have substantial impacts without generating important expectations–formation effects;
- a small but persistent intervention is more likely to destabilize a linear model than is a large but fleeting intervention—a new finding that refines a point that Lucas (1976) emphasizes.

We apply the approach to policy analysis developed in the theory to an empirical model fit to post-war U.S. data. The model projects the macroeconomy conditional on hypothetical policies designed to address questions Federal Reserve officials may have asked. In the identified vector autoregression (VAR) we estimate, the conditional projections reflect only the direct effects of policy. For every policy, we report the modesty statistic to determine whether the direct effects are large relative to their historical variation. Whenever direct effects are unusually large—which may or may not trigger substantive shifts in beliefs—we label it *immodest* and we infer that the VAR is likely to perform badly by missing important expectations–formation effects.

We display several examples from the U.S. experience. In some cases the modesty statistic calls for healthy skepticism about the VAR. In other cases one can feel confident that the policy change, if implemented, would not generate substantial expectations–formation effects. Many intra-regime policy interventions that might arise routinely in Federal Open Market Committee (FOMC) discussions are modest relative to the Lucas critique, but can shift the projected paths and probability distributions of macro variables in ways that matter to policymakers' decisions.

2. Contacts with literature

This work makes several points of contact with existing literature. To some readers, the theoretical environment described in the introduction may resemble the one in which the Federal Reserve operates. With no mechanism binding the Fed's choices in the future, a skeptical public bases its beliefs more on the actions of the FOMC than on the statements of Fed officials. This skepticism makes the history of actions matter. Although many observers credit the decline of inflation in the 1980s to a shift in monetary policy behavior, even the most sanguine proponents of recent Fed successes cannot exclude the possibility of a return to inflationary policies, as in Sargent's (1999) analysis of American inflation.

Despite the description's ring of relevance, macropolicy analyses typically do not use this environment. Standard analyses instead mimic the examples Lucas (1976) used to illustrate his critique of econometric policy evaluation. Rather than an ongoing process, in Lucas's thought experiment policy choice is once-and-for-all.

Several authors have pointed out a problem with posing policy as making once-and-for-all choices: it is logically inconsistent (e.g., Cooley et al., 1982, 1984; Sargent, 1984, Sims, 1982, 1987, 1998). Treating regime changes as surprises that will never occur again ascribes to the public beliefs about policy that are inconsistent with actual behavior—the government takes actions that the public thought were impossible. Cooley et al. (1982, p. 2) resolve the inconsistency by proposing the general principle for policy analysis that "…any entity which changes over time in a way that is not completely predictable should be modeled as a sequence of random variables." In applying this principle, Cooley et al. lay a probability distribution over all possible rules and define a policy intervention as a sequence of realizations of policy variables, rather than a permanently new rule. Agents' decision rules then incorporate the belief that it is always possible to return to the good (bad) old days of policymaking.

Viewing policy regime as evolving stochastically leads to an interpretation of the Lucas critique that expands on Sims's (1998) observation that Lucas can be interpreted as pointing towards a potentially important source of nonlinearity. In the face of sustained dynamic patterns of change in policy variables, agents may grow to believe that policy behavior has shifted, and respond by adjusting their expectations and decision rules accordingly. This interpretation does not diminish the potential force of Lucas's critique; instead it provides a framework for isolating and quantifying the effects of the behavior that Lucas emphasizes.

Finally, in using a loosely identified VAR to predict the consequences of counterfactual policy scenarios, we draw on Hurwicz's (1962) insight that the weaker are a model's identifying assumptions, the smaller is the class of modifications for which it is "structural," or invariant. Our VAR identifies only money demand and monetary policy behavior, leaving other aspects of the macroeconomy in reduced form. This identification scheme necessarily limits the class of policy interventions for which the VAR's predictions will be reliable. We show nonetheless that this class includes many policy scenarios that the Federal Reserve routinely contemplates.

3. A theoretical framework for policy analysis

The theoretical framework precisely defines a modest policy intervention. Examples show cases where an intervention generates large direct effects without inducing the shifts in agents' beliefs about regime that create substantial expectations–formation effects. Examples also display interventions whose direct and expectations–formation effects are both large. In the theory, policy analysis consists of forecasts conditional on interventions. This corresponds to procedures the Federal Reserve employs and to the empirical exercises Section 5 implements.

3.1. The model

The model extends Cochrane's (1998) setup to allow policy behavior to switch randomly between two regimes. Price-setting behavior is based on the costly price adjustment framework of Rotemberg (1982, 1996), and generates predictions that are consistent with models that introduce nominal rigidities through other means. Cochrane combines a simple aggregate demand specification with the dynamic price-setting behavior. Costly price adjustment allows expected monetary policy to have both price and output effects.

Pricing behavior involves inertial and forward-looking components. Costs of adjusting prices are parameterized by $\alpha \in [0, 1]$, which determines the degree of stickiness. A representative monopolistically competitive firm chooses the price sequence $\{p_t\}$ to maximize profits discounted at rate $\beta \in (0, 1)$, leading to the price-adjustment equation:

$$p_t = \alpha p_{t-1} + (1-\alpha)(1-\alpha\beta)\mathrm{E}_{t-1}\sum_{j=0}^{\infty}(\alpha\beta)^j m_{t+j}, \tag{1}$$

where m is the nominal money stock and all variables are in logarithms. A simple money demand expression links output and money: $y_t = m_t - p_t$.

Combining (1) with the demand for money implies the expression for equilibrium output

$$y_t = \left[m_t - \frac{1-\alpha}{1-\alpha L}\mathrm{E}_{t-1}\frac{1-\alpha\beta}{1-\alpha\beta L^{-1}}m_t\right]. \tag{2}$$

As Cochrane observes, driving adjustment costs to zero ($\alpha \to 0$) reduces the model to Lucas's (1972, 1973) environment in which only unanticipated money affects output: $y_t = m_t - \mathrm{E}_{t-1}m_t$. As $\alpha \to 1$, adjustment costs become infinite and the distinction between anticipated and unanticipated policy disappears: $y_t = m_t - p$, where p is the constant price level.

To complete the model we posit a process for setting the money supply that randomly switches between two policy regimes. Let g_t be the growth rate of the nominal money supply between $t-1$ and t:

$$m_t = g_t + m_{t-1}, \tag{3}$$

given some initial money stock, $m_0 > 0$. Denoting the regime at t by R_t, the policy rule is

$$g_t = \mu(R_t) + \rho(R_t)g_{t-1} + \sigma(R_t)\varepsilon_{Pt}, \quad \varepsilon_{Pt} \sim \mathrm{N}(0,1), \quad g_0 > 0. \tag{4}$$

The specification allows for the possibility that a shift in policy regime changes the dynamics, the mean, and the variance of money growth. ε_P is an intra-regime exogenous shift in policy.

Label the two policy regimes R^1 and R^2.[2] Regime switches obey a Markov chain with transition probabilities

$$\mathrm{P} = \begin{bmatrix} \mathrm{P}[R_t = R^1 | R_{t-1} = R^1] & \mathrm{P}[R_t = R^1 | R_{t-1} = R^2] \\ \mathrm{P}[R_t = R^2 | R_{t-1} = R^1] & \mathrm{P}[R_t = R^2 | R_{t-1} = R^2] \end{bmatrix} = \begin{bmatrix} p_{11} & 1 - p_{22} \\ 1 - p_{11} & p_{22} \end{bmatrix}, \tag{5}$$

and associated policy parameters

$$(\mu(R_t), \rho(R_t), \sigma(R_t)) = \begin{cases} (\mu_1, \rho_1, \sigma_1) & \text{if } R_t = R^1, \\ (\mu_2, \rho_2, \sigma_2) & \text{if } R_t = R^2. \end{cases} \tag{6}$$

The *policy process* is completely defined by Eqs. (3)–(6) and values for the vector of policy parameters $\Pi \equiv (\mu_1, \mu_2, \rho_1, \rho_2, \sigma_1, \sigma_2, p_{11}, p_{22})$. A *realization of policy* at t is the pair (g_t, R_t). Let $\Omega_t = \{p(R_0), m_0, g_0, g_1, \ldots, g_t\}$ where $p(R_0)$ is agents' prior belief about regime at the time 0. Agents' decisions at t are based on information contained in the set Ω_{t-1}, along with Π and their updated beliefs about regime, $\mathrm{P}(R_{t-1} = R^s | \Omega_{t-1})$, for $s = 1, 2$. We assume agents observe the history of money growth realizations, but none of the realizations of regime.

3.2. Direct and expectations–formation effects

When regime is fixed, the model is linear and has a vector autoregressive representation. The K-period forecast from the VAR under Regime 1, given information at T, is

$$x_{T+K} = \sum_{s=0}^{K-1} C_s \varepsilon_{T+K-s} + \mathrm{E}(x_{T+K} | \Omega_T, R_t = R^1, t = T+1, \ldots, T+K), \tag{7}$$

where $x_t = (p_t, y_t, m_t)'$ is a vector of variables from the model, C_s is the impulse response matrix at horizon s, and $\varepsilon_{T+K-s} = (0, 0, \varepsilon_{P,T+K-s})'$. Direct effects and expectations–formation effects from a policy intervention are computed from forecasts from fixed-regime and switching-regime models conditional on a given intervention. Both of these forecasts are reported relative to a baseline forecast denoted by $\mathrm{E}(x_{T+K} | \Omega_T, R_t = R^1, t = T+1, \ldots, T+K)$ in (7).

Let I_T be a hypothetical intervention at time T, specified as a K-period sequence of exogenous policy actions, $I_T = \{\tilde{\varepsilon}_{P,T+1}, \ldots, \tilde{\varepsilon}_{P,T+K}\}$. Although the policy advisor chooses I_T, private agents treat it as random. Denote a forecast of $\{x_{T+K}\}$ conditional on I_T by $\mathrm{E}(x_{T+K} | I_T, \Omega_T, R_t = R^1, t = T+1, \ldots, T+K)$. Direct effects

[2] For analytical clarity, we consider only two regimes. Our analysis can be easily extended to three or more regimes. No conclusions or logical points hinge on the two-regime assumption. Cooley et al. (1982) and Andolfatto and Gomme (2003) consider less general regime-switching specifications in general equilibrium models.

of I_T relative to baseline are

$$\begin{aligned}\text{Direct effects} \equiv \eta_{P,T+K} &= \sum_{s=0}^{K-1} C_s \tilde{\varepsilon}_{P,T+K-s} \\ &= \mathrm{E}(x_{T+K}|I_T, \Omega_T, R_t = R^1, t = T+1, \ldots, T+K) \\ &\quad - \mathrm{E}(x_{T+K}|\Omega_T, R_t = R^1, t = T+1, \ldots, T+K). \end{aligned} \tag{8}$$

The second conditional expectations term—the baseline forecast—is equivalent to an intervention that sets $\tilde{\varepsilon}_{P,s} = 0,\ s = T+1, \ldots, T+K$.

When regime may change, the economy is nonlinear and the economy no longer takes the VAR form in (7). Now an intervention may trigger changes in agents' beliefs about policy regime, which affect agents' expectations of future policy and their optimal choices. *Total effects* of policy combine direct effects with *expectations–formation effects*, induced by changes in beliefs about regime. Total effects relative to the baseline projection in the linear model are:

$$\begin{aligned}\text{Total effects} \equiv \mathrm{E}(x_{T+K}|I_T, \Omega_T) - \mathrm{E}(x_{T+K}|\Omega_T, R_t = R^1, \\ t = T+1, \ldots, T+K), \end{aligned} \tag{9}$$

where the same intervention, I_T, is conditioned on in (8) and (9).

With direct effects and total effects computed relative to the same baseline, we isolate expectations–formation effects, defined as the difference between (9) and (8):

$$\text{Expectations}-\text{formation effects} \equiv \text{total effects} - \text{direct effects}. \tag{10}$$

Expectations–formation effects arise from the changes in behavior that lie at the heart of Lucas's (1976) critique. A natural way to judge whether the Lucas critique is quantitatively important is to check if expectations–formation effects are small. If those effects are small, then forecasts from a model that assumes policy regime is fixed will be reasonably accurate. If, in contrast, expectations–formation effects are large relative to the standard variation in direct effects,[3] then the fixed-regime model's predictions will be systematically wrong. This is the situation on which Lucas focuses.

3.3. Simulating the model

The theory offers a laboratory for finding examples of interventions where the Lucas critique bites. By separating direct and expectations–formation effects, the theory implies a natural measure of whether a particular intervention is modest (Section 3.3.1).

Of course any inferences about whether the Lucas critique bites for a given intervention depend on the model's parameterization. We focus on two different sets of parameters. In the first set, policy regimes are far apart and shifts in beliefs about

[3] It is the size of expectations–formation effects relative to the *historical variation* of direct effects—not relative to the direct effect itself—that matters for policy analysis. If both direct and expectations–formation effects are small, for example, the ratio of the two is meaningless.

regime can generate quantitatively important expectations–formation effects under certain conditions (Section 3.3.2). The second set makes the two regimes much closer and expectations–formation effects tend to be small for many interventions (Section 3.3.3).

To simulate the model we specify values for the private parameters, (α, β), and the policy parameter vector, Π. Taking the period of the model to be monthly, as in the identified VAR estimated in Section 4, we set $\beta = 0.99674$, which yields an average annual real interest rate of 4.0 percent. To make expectations–formation effects potentially large, we set $\alpha = 0.9$, implying costly price adjustment and important impacts of expected policy on output.[4] Transition probabilities are chosen as $p_{11} = 0.99722$, $p_{22} = 0.99167$, so that in a monthly model Regime 1 (low money growth) lasts 30 years on average, and Regime 2 (high money growth) lasts 10 years.

3.3.1. *A statistic for modest interventions*

We maintain the assumption that policy has been operating under Regime 1 and that agents' beliefs in this regime are firmly held. For an intra-regime intervention, the distribution of direct effects may be obtained from the sequence of forecast errors computed in (8). The statistic $\eta^*_{P,T+K}$ for measuring a direct effect, which scales $\eta_{P,T+K}$ by the standard variation of the direct effect $\sqrt{\sum_{s=0}^{K-1} C_s^2}$, has a standard normal distribution. For comparison, total effects and expectations–formation effects are also measured relative to the standard variation of direct effects.

Definition. An intervention is *modest* if over a specified forecast horizon, K, and for variable i, $|e_i \eta^*_{PT+K}|$ is not close to 2 (standard deviations), where e_i is a row vector of zeros with unity in the ith column.[5]

The modesty statistic η^* reports how unusual a conditional projection is relative to the typical size of the direct effects, as measured in units of standard deviations of direct effects. Here we take a probabilistic point of view (Sims and Zha, 1999): we want to gauge how large the direct effects of a hypothetical intervention are and how likely the intervention is to trigger changes in agents' behavior. When direct effects fall within their normal variation, as judged by two standard deviations, the effects are typical under Regime 1. When direct effects are typical, there is no reason for agents to change their beliefs about the prevailing regime, and expectations–formation effects are likely to be insignificant.

The likelihood of regime change implied by the modesty statistic is important. If two policy regimes are very close in the sense that typical direct effects under Regime 2 are well within two standard deviations of direct effects under Regime 1, these two regimes, although different analytically, are empirically indistinguishable. Thus, our statistic offers a necessary but not sufficient condition for the Lucas critique to bite quantitatively. An example of this is discussed in Section 3.3.3.

[4] When $\alpha = 0$ both direct and expectations–formation effects on output and expectations–formation effects on prices are small.

[5] The statistic is in the spirit of Doan et al. (1984) "implausibility index."

When the two regimes are far apart, there is a close relationship between large values of the modesty statistic and large expectations–formation effects. In this case, learning about policy regime plays a central role. With policy regime being a hidden state variable, agents behave rationally as Bayesian updaters to use the observed information set to infer the probability of policy being in each of the two regimes. Suppose agents have initial beliefs about R_{t-1} denoted by $\mathrm{P}(R_{t-1} = R^s|\Omega_{t-1})$ for $s = 1, 2$. At time $t+h$, when Ω_{t+h} is available, agents compute $\mathrm{P}(R_{t+h}|\Omega_{t+h})$ recursively as

$$\mathrm{P}(R_{t+h}|\Omega_{t+h}) = \frac{\varphi(g_{t+h} - \mu(R_{t+h}) - \rho(R_{t+h})g_{t+h-1}; \sigma^2(R_{t+h}))\mathrm{P}(R_{t+h}|\Omega_{t+h-1})}{\sum_{R_{t+h}=R^1,R^2}\{\varphi(g_{t+h} - \mu(R_{t+h}) - \rho(R_{t+h})g_{t+h-1}; \sigma^2(R_{t+h}))\mathrm{P}(R_{t+h}|\Omega_{t+h-1})\}}, \tag{11}$$

where $\varphi(x; y)$ is the normal probability density with mean x and variance y, and

$$\mathrm{P}(R_{t+h}|\Omega_{t+h-1}) = \sum_{R_{t+h-1}=R^1,R^2}\{\mathrm{P}(R_{t+h}|R_{t+h-1})\mathrm{P}(R_{t+h-1}|\Omega_{t+h-1})\}. \tag{12}$$

The posterior probability in (11) becomes the agents' prior beliefs in the next period. The relative likelihood weight that appears in the updated beliefs is

$$\frac{\varphi(g_{t+h} - \mu(R_{t+h}) - \rho(R_{t+h})g_{t+h-1}; \sigma^2(R_{t+h}))}{\sum_{R_{t+h}=R^1,R^2}\varphi(g_{t+h} - \mu(R_{t+h}) - \rho(R_{t+h})g_{t+h-1}; \sigma^2(R_{t+h}))}.$$

The likelihood weight is an important factor in determining how agents' beliefs change in response to new information about policy. Suppose that at time $t-1$ agents hold a strong belief that the current regime is 1. Imagine that intra-regime interventions at time t alter the relative likelihood weight in favor of Regime 2 so that the weight is

$$w_2 = \frac{\varphi(g_t - \mu_2 - \rho_2 g_{t-1}; \sigma_2^2)}{\sum_{R_t=R^1,R^2}\varphi(g_t - \mu(R_t) - \rho(R_t)g_{t-1}; \sigma^2(R_t))} > 0.5. \tag{13}$$

Then beliefs shift toward Regime 2. It follows from (13) that the larger is the modesty statistic, the larger is the weight w_2, and the faster beliefs change in favor of Regime 2. If w_2 were near 1, agents would feel nearly certain that Regime 2 prevails and expectations–formation effects may be substantial. Section 3.3.2 shows some examples.

3.3.2. *Policy regimes with extreme differences*

In this and the next section we consider two kinds of interventions (labeled A and B) and some variations (C and D). The sections illustrate that small but persistent interventions are likely to trigger important changes in agents' belief, while large but short-lived ones, contrary to the conventional view, tend not to trigger changes in beliefs that have quantitatively important implications. We choose a 4-year forecast horizon to coincide with a typical horizon the Federal Reserve considers in their routine policy making and with the empirical exercises conducted in Section 5.

We begin with regimes that are far apart. For the extreme differences in regime, we choose rules that imply average annual money growth rates of 3.04 and 13.08 percent, slightly different serial correlation patterns, and the same variance of the policy shock across regimes. The policy regimes are:

$$\begin{aligned} &\text{Regime 1: } g_t = 0.0005 + 0.800 g_{t-1} + 0.0015\varepsilon_{Pt}, \\ &\text{Regime 2: } g_t = 0.0007 + 0.932 g_{t-1} + 0.0015\varepsilon_{Pt}. \end{aligned} \tag{14}$$

Consider first the special case of the one-period intervention described by $I_T = \{1, 0, \ldots, 0\}$. By conditioning on the counterfactual that Regime 1 prevails now and in the future, the exercise corresponds to conventional impulse response analysis. This intra-regime monetary expansion generates a hump-shaped increase in output that peaks at over 0.3 percent after seven months, and smooth increases in both prices and the money stock that converge to permanently higher (and equal) levels. Monetary surprises have their biggest impacts on output and inflation within a year. The hump-shaped path of output and the smooth, gradual response of prices resemble the policy impacts found in the empirical model of section, though the timing and magnitudes differ. Direct effects of alternative policy interventions are functions of this conventional impulse response.

Specification A. We turn now to an intervention that is sustained but relatively small (less than one standard deviation) in each of 48 months: $I_T = \{2/3, 2/3, \ldots, 2/3\}$.[6] The first panel in Fig. 1 records the percentage deviation from the baseline forecast of the money stock and the posterior probability that policy is in Regime 1. Direct and expectations–formation effects on prices and output appear in the next two panels on the left side of the figure, expressed as percentage changes from baseline. These are cumulative impacts: in month j the dynamic impacts of all the interventions from $T+1$ to $T+j$ appear. The intervention persistently puts higher likelihood weight on Regime 2, smoothly shifting agents' beliefs away from Regime 1. As beliefs evolve, expectations–formation effects grow in importance. Because price adjustment is costly ($\alpha = 0.9$), the expectations–formation effects show up more strongly in output than in prices. Direct effects and expectations–formation effects on prices are reinforcing, though they have opposite influences on output. At 48 months the total effects on output are close to zero: the cumulative expectations–formation effects reduce output by nearly 4 percent, offsetting the increase due to direct effects.

Table 1 summarizes these impacts and reports the modesty statistic for this intervention in the 48th month (Specification A under extreme case). Absolute values of η^* close to 2 or larger imply the intervention is likely to be immodest. The table also reports expectations–formation effects. Although expectations–formation effects are large only in the case of output, the modesty statistic implies that the resulting paths of both output and prices are unlikely to arise in the absence of

[6] The linear model in Regime 1 requires a sequence of $\tilde{\varepsilon}_P$'s on the order of 1.0 to mimic policy behavior in Regime 2 with zero intra-regime shocks. Although shocks of 2/3 do not put policy realizations at the center of Regime 2, they are sufficient to change agents' beliefs. Unless otherwise noted, initial beliefs about regime were set at $P(R_T = R^1) = 0.98$ in Section 3.

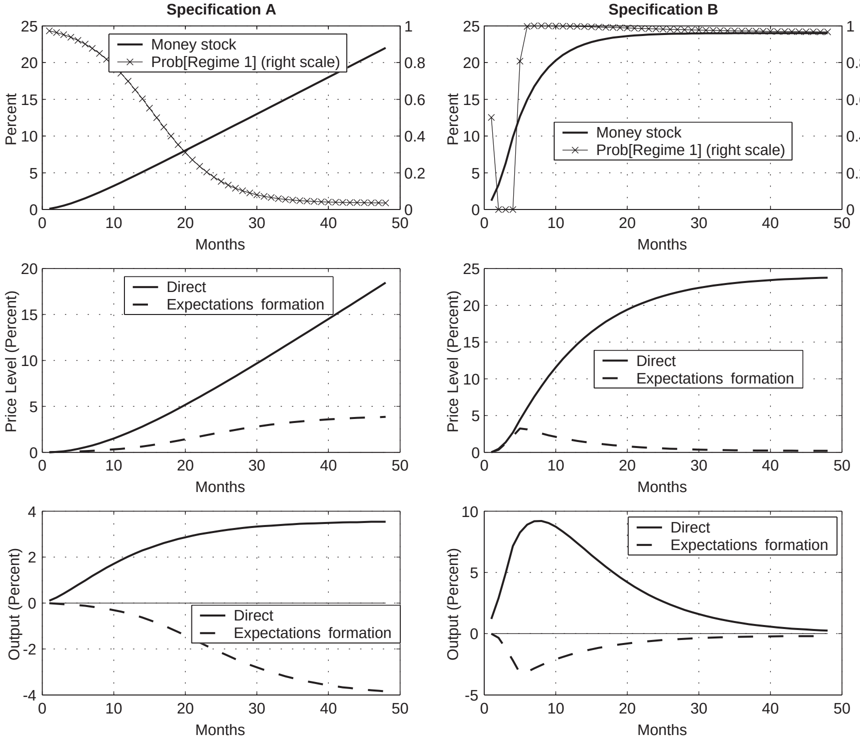

Fig. 1. Impacts of two hypothetical interventions (extreme differences in policy regimes).

substantial expectations–formation effects. This makes clear the sense in which our statistic is necessary for signaling a linear model will breakdown.

Specification B. We now alter the dynamic pattern of the intervention to consider an intervention that is large but short-lived: $I_T = \{8, 8, 8, 8, 0, \ldots, 0\}$. The cumulative size of the intervention is the same as in Specification A. This intervention produces extremely rapid money growth for the first four months. The money growth rates caused by these huge shocks are so unlikely to occur under Regime 1 that agents quickly become nearly certain policy switched to Regime 2 (the first panel on the right side of Fig. 1).[7] With a lack of persistence in the intervention, as soon as the shocks disappear agents revert to believing Regime 1 prevails. The direct and expectations–formation effects on output are shrunken mirror images of each other and closely resemble the shape of the underlying impulse response functions (the last two panels on the right side of the figure). Over very short horizons, direct effects and expectations–formation effects on the price level are nearly identical, but direct effects soon become dominant at longer horizons. The cumulative

[7] If there were a third regime that mimics these policy interventions, beliefs would switch to that regime. This complication does not change the key point that for a short time agents move *away* from believing that Regime 1 prevails. Where their beliefs shift *to* is not central.

Table 1
Impacts of policy interventions at 4-year horizon

Specification	Direct effects $\eta^*_{P,T+K}$ (Standard deviations) p	y	Expectations–formation effects (Standard deviations) p	y
Extreme case				
A	4.53	3.45	0.86	−3.50
B	5.34	0.22	0.04	−0.17
C	2.07	1.61	0.16	−0.63
D	4.53	3.45	0.30	−1.20
Less extreme case				
A	4.54	3.32	0.02	−0.06
B	5.55	0.19	0.01	−0.05

Direct effects (η^*) and expectations–formation effects scaled by standard errors of direct effects based on 5000 draws. Interventions: A: $\tilde{\varepsilon}_P = 0.667$ in each of 48 months. B: $\tilde{\varepsilon}_P = 8.0$ for first 4 months, $\tilde{\varepsilon} = 0$ for next 44 months. C: $\tilde{\varepsilon}_P = 0.333$ in each of 48 months. D: $\tilde{\varepsilon}_P = 0.667$ in each of 48 months, but $p_{22} = 0.9167$ (1-year duration of Regime 2). In specifications A–D, $P(R_T = R^1) = 0.98$. "Less extreme case" calibrated to U.S. data by splitting sample into two "regimes" 1959:2–1971:12/1983:4–2000:07 and 1972:1–1982:12 (excluding 1983:1–1983:3 due to exceptionally high money growth rates) and fitting AR (1) processes to the monthly growth rate of $M2$ in each period.

expectations–formation effects at the 48th month are very small on both variables. This outcome is confirmed by the modesty statistic for output (Table 1).[8] This is a situation where the horizon of concern plays an important role in determining the usefulness of the linear model. Policymakers concerned with very short horizons may choose to down weight the model's influence. For policymakers with horizons of two or more years, the message of the simulation is that a short-lived intervention—even a large intervention—is likely to generate only small expectations–formation effects over the relevant horizons. At those horizons the predictions of a linear econometric model are likely to be reliable.

Specification C. Table 1 also reports how sensitive results are to the total size of the intervention. Specification C cuts in half the magnitude of the intervention in Specification A, but maintains the sustained nature of the intervention. In this case, the expectations–formation effects on both output and prices are small and the modesty statistic raises no warning flags for output, though the statistic slightly exceeds 2 for prices. Even a sustained intervention may not undermine the accuracy of a linear approximation if the intervention is sufficiently small.

Specification D. Regime duration influences agents' beliefs. Specification D reproduces Specification A but reduces the duration of Regime 2 from 10 to 1 year. Making Regime 2 ephemeral greatly increases the likelihood that Regime 1 produced the realized path of money growth; agents' posterior probability of Regime 1 falls

[8] The large statistic for the price level is another example where an intervention can have large direct effects without generating significant expectations–formation effects.

only to 0.38 (as compared to 0.04 in Specification A). As consequence, expectations–formation effects on output, which were substantial when Regime 2 was longer lived, are now small. Because the direct effects of the intervention on output and prices are identical to Specification A, the modesty statistic is unchanged.[9]

3.3.3. Policy regimes with less extreme differences

Extreme differences in regimes can translate changes in beliefs about regime into substantial expectations–formation effects. When differences in regimes are less pronounced, shifts in beliefs need not generate important expectations–formation effects. This section reports sensitivity analysis by considering regimes that are closer than those in Section 3.3.2.

We loosely calibrate the policy process to U.S. data on monthly $M2$ growth rates since 1959.[10] The calibration breaks the sample into two "regimes" that imply average annual money growth rates of 6.4 percent in Regime 1 and 9.4 percent in Regime 2; the lower money growth regime is associated with somewhat greater persistence and lower variance of the shock. The regimes are:

$$\begin{aligned} &\text{Regime 1: } g_t = 0.0013 + 0.75 g_{t-1} + 0.0019 \varepsilon_{Pt}, \\ &\text{Regime 2: } g_t = 0.003 + 0.60 g_{t-1} + 0.0024 \varepsilon_{Pt}. \end{aligned} \tag{15}$$

When regimes exhibit less extreme differences the same interventions considered above yield dramatically different results. Consider again the sustained interventions of 2/3 of a standard deviation each period, Specification A. The intervention generates substantial direct effects, raising the price level 19 percent and output 4 percent above their baseline levels. Agents' beliefs about regime do not move from their initial belief that Regime 1 is very likely. The reason is that the conditional likelihood under Regime 2 is more dispersed than under Regime 1 ($\sigma_2 > \sigma_1$), so the interventions have to be larger to make Regime 2 more likely. Reinforcing the influence of the relative variances of shocks in the two regimes is that money growth is less persistent in Regime 2 ($\rho_1 > \rho_2$), similar to Specification D in Section 3.3.2. If realized money growth exceeds the unconditional mean in Regime 2, agents would expect more rapid mean reversion in Regime 2 than is consistent with the persistent intervention. This makes Regime 2's conditional likelihood value less than Regime 1's. Unchanged beliefs lead to very small expectations–formation effects in spite of the substantial direct effects the intervention produces. The modesty statistic indicates direct effects on a par with those under the extreme specification of regime differences (Table 1). This example underscores our central message that expectations–formation effects can be tiny even though direct effects are large.

The large, but brief, intervention in Specification B in Table 1 generates rapid changes in beliefs about regime, but miniscule expectations–formation effects. When

[9] If the maintained assumption that agents believe firmly in the prevailing regime does not hold, it is possible for the modesty statistic η^* to mislead. In special circumstances an intervention may produce large enough expectations–formation effects to cause trouble for the linear model, yet generate direct effects small enough for the intervention to be deemed modest. Leeper and Zha (2002b) discuss this situation and propose a method for detecting when it may arise.

[10] Table 1 describes the calibration.

the large interventions end, agents return to believing that Regime 1 prevails. Expectations–formation effects on both prices and output are very small, even in the short run, despite large direct effects that raise output by more than 10 percent in the short run. By 48 months, the cumulative direct effects on prices exceed those under Specification A, but the direct effects on output have died out.

In the empirical analysis below, the interventions we consider that correspond to actual Federal Reserve behavior are smaller and briefer than those studied in the theoretical examples. The empirical interventions are modest according to our statistic. Although they do not run afoul of the Lucas critique, they generate economically meaningful direct effects.

4. The econometric model

This section specifies and estimates a small structural model of American monetary policy behavior. The section reports the model's impulse response functions, which are the direct effects of a particular policy intervention, and discusses the model's fit, stability, and suitability for policy analysis.

4.1. The model

Actual policy behavior is a complicated function of a high-dimensional vector of variables. Policymakers choose R_t, the vector of policy choices at date t, as a function of their information set, Ω_t. Actual policy behavior is a function g such that

$$R_t = g(\Omega_t). \tag{16}$$

We assume that private agents are not privy to the details of the policymakers' decision problems, including the policymakers' incentives and constraints. That is, they observe the information set $S_t \subset \Omega_t$. Agents perceive that policy is composed of a regular response to the state of the economy that they observe at time t, S_t, and a random part, ε_{Pt}. The econometric model of policy is

$$R_t = f(S_t) + \varepsilon_{Pt}. \tag{17}$$

We take f to be linear. ε_{Pt} is exogenous to the econometric model. A policy regime is a choice of g, which implies an f and a stochastic process for ε_P.

The econometric model embeds the policy behavior in (17) in a system of equations. If y_t is an $(m \times 1)$ vector of time series, the structural form is

$$\sum_{s=0}^{p} A_s y_{t-s} = \varepsilon_t, \tag{18}$$

where ε_t is a vector of i.i.d. structural disturbances that are exogenous to the model. Those disturbances hit both nonpolicy and policy sectors of the economy, so

$$\varepsilon_t = \begin{bmatrix} \varepsilon_{Nt} \\ \varepsilon_{Pt} \end{bmatrix}, \tag{19}$$

where ε_{Nt} is the vector of nonpolicy disturbances.

The environment developed in Section 3 addresses policy questions by using a structure like (18) to project y conditional on hypothesized paths for ε_P (Marschak, 1953). We assume ε_{Pt} is a vector of exogenous random variables, uncorrelated with all the nonpolicy exogenous disturbances in the economy. The errors are Gaussian with

$$\mathrm{E}(\varepsilon_t \varepsilon_t' | y_{t-s}, s > 0) = I, \quad \mathrm{E}(\varepsilon_t | y_{t-s}, s > 0) = 0 \quad \text{all } t. \tag{20}$$

The A_s matrices and the probability distribution of ε define the model's structure.

Assuming the matrix of contemporaneous coefficients, A_0, is nonsingular, there is a representation of y in terms of the impulse responses functions:

$$y_t = \sum_{s=0}^{t-1} C_s \varepsilon_{t-s} + \mathrm{E}_0 y_t. \tag{21}$$

Elements of C_s report how each variable in y responds over time to the behavioral disturbances in ε. $\mathrm{E}_0 y_t$ is the projection of y_t conditional on initial conditions. The reduced form of (18) is

$$\sum_{s=0}^{p} B_s y_{t-s} = u_t, \tag{22}$$

with $B_0 = I$ and the covariance of the reduced-form errors, u, is $\Sigma = A_0^{-1} A_0^{-1'}$.

Expressions (18) and (22) imply a linear mapping from the reduced-form errors to the behavioral disturbances:

$$u_t = A_0^{-1} \varepsilon_t. \tag{23}$$

Identification of the structural form follows from imposing sufficient restrictions on A_0 so that there are no more than $m(m-1)/2$ free parameters in A_0.

4.2. Identification

We estimate a model that contains six variables and three sectors. The identification scheme follows the general approach in Gordon and Leeper (1994) and Sims and Zha (1998). There are no restrictions on lagged variables. Three goods market variables—real GDP (y), consumer prices (P), and the unemployment rate (U)—compose the production sector, and are the ultimate objectives of monetary policy. We do not model the markets for reserves and a broad monetary aggregate, opting for compactness to treat the $M2$ money stock as the aggregate and the federal funds rate (R^f)—the monetary policy instrument—as the price that clears the money market.

The third sector describes an "information variable"—commodity prices (CP)—that is available at high frequencies and reacts instantaneously to shocks from all sectors of the economy. The data, which the appendix describes, are monthly from January 1959 to September 1998. Monthly GDP is interpolated from quarterly GDP. All data are logarithmic except for the federal funds rate and the unemployment rate. We estimate the model with 13 lags.

The identification treats the production sector as predetermined for the rest of the system, reflecting the view that production, pricing, and employment decisions do not respond immediately to shocks from outside the sector. Production sector variables interact only with each other within the period. Money market variables and information variables do not enter this sector, reflecting sluggishness in the goods market due to contracts and advance planning of production. Distinct behavioral equations within the production sector are not identified; instead, the coefficients are arranged in lower triangular form in the order y, P, and U.

The monetary sector allows simultaneity in determining the money stock and the federal funds rate. The demand for nominal money balances depends on the short-term nominal interest rate, real income (proxied by real GDP), and the price level. We do not impose short-run homogeneity in prices.

We base the specification of monetary policy behavior in (17) on the information available to the Federal Reserve within the month. During the month, the Federal Reserve sets its interest rate instrument based on current observations on the money stock and commodity prices.[11] We set to zero the coefficients on y, P, and U in the policy function because within the month the Federal Reserve does not observe these variables directly. The error term in the monetary policy equation, ε_P, represents intra-regime exogenous policy interventions.

4.3. Parameter estimates and direct effects of policy

For the two behavioral equations we report estimated contemporaneous coefficients along with 68 percent equal-tailed probability intervals for the behavioral coefficients, estimated over the full sample period.[12] Those equations are money demand:

$$\underset{(49.65,\ 434.46)}{310.33M2} + \underset{(45.33,\ 184.55)}{161.89R^f} - \underset{(-36.44,\ -9.11)}{28.47y} + \underset{(-23.95,\ 36.95)}{6.84P} = \varepsilon_{\mathrm{MD}} \tag{24}$$

and monetary policy:

$$\underset{(-12.08,\ 169.12)}{100.66R^f} - \underset{(-444.63,\ -40.42)}{336.84M2} - \underset{(-11.29,\ 4.90)}{3.10\mathrm{CP}} = \varepsilon_P\,. \tag{25}$$

Money demand has reasonable economic interpretations: the interest elasticity of demand is negative; the output elasticity is positive; the price elasticity is small and

[11] We allow policy to choose not to react strongly to commodity prices by shrinking the prior standard deviation on the coefficient of CP toward the zero prior mean by a factor of 0.05.

[12] Probability intervals are reported in parentheses. Those intervals are exact finite-sample results computed by a Gibbs sampler algorithm with 360,000 Monte Carlo draws (Waggoner and Zha, 2002, 2003). Appendix A in Leeper and Zha (2002a) describes the Bayesian estimation methods employed.

imprecisely estimated. Monetary policy responds strongly to the money stock: disturbances that raise the money stock induce the Fed to increase the federal funds rate. Although the estimates seem to suggest the Fed does not react much to information contained in commodity prices, this interpretation may be misleading. In the policy equation the coefficients on R^f and CP are highly correlated (0.97), as are the coefficients on $M2$ and CP (0.72); these correlations muddy inferences about individual coefficients.

We base a probabilistic assessment of the model's overall fit on the shape of the likelihood surface, not on tests of whether individual coefficients are different from zero (Sims and Zha, 1999). But the probability intervals for individual parameters reported in Eqs. (24) and (25) show that they are skewed, with most of the probability mass concentrated around the maximum likelihood estimates. Because we care about the equilibrium effects of exogenous policy actions, none of our inferences rest on an individual parameter in A_0.

Impulse response functions from an estimated VAR correspond to the direct effects we describe in Section 3.2 for a fixed-regime version of the theoretical model. Fig. 2 displays the C_s's in (21) over 48 months for an exogenous monetary policy contraction. The solid lines are the maximum likelihood estimates of responses and the dashed lines are equal-tailed error bands containing 0.68 probability.

The contraction raises the funds rate initially and immediately decreases the money stock and commodity prices, both of which continue to decline smoothly over the four-year horizon. After a brief delay, output falls and stays lower, while unemployment rises. Six months after the exogenous action, both output and unemployment are likely to differ from their initial levels. Consumer prices adjust more slowly and are unlikely to be appreciably lower for about a year. After a year prices decline smoothly and remain well below their initial level.

The response of the interest rate to an exogenous policy contraction stands out. In Fig. 2 the initial liquidity effect lasts about eight months, but by 18 months the funds rate lies well below its initial level. The lower funds rate then persists. This is the shape of the path of the short-term nominal interest rate following a monetary contraction that Friedman (1968) and Cagan (1972) describe as a short-lived liquidity effect followed by income and expected inflation effects. After four years the declines in inflation and the federal funds rate are the same size, as one might anticipate if expected inflation is the dominant source of fluctuations in nominal rates over long periods. The responses in the figures suggest that to lower inflation persistently the Fed should raise the funds rate only briefly. Because lower inflation is ultimately associated with a lower funds rate, the Fed must begin to reduce the rate within about a year, and then keep it lower.

4.4. Suitability for policy analysis

Any model to be used for policy analysis must be carefully scrutinized. Although our model is overidentified, the restrictions do not invalidate the uses to which we put the model.

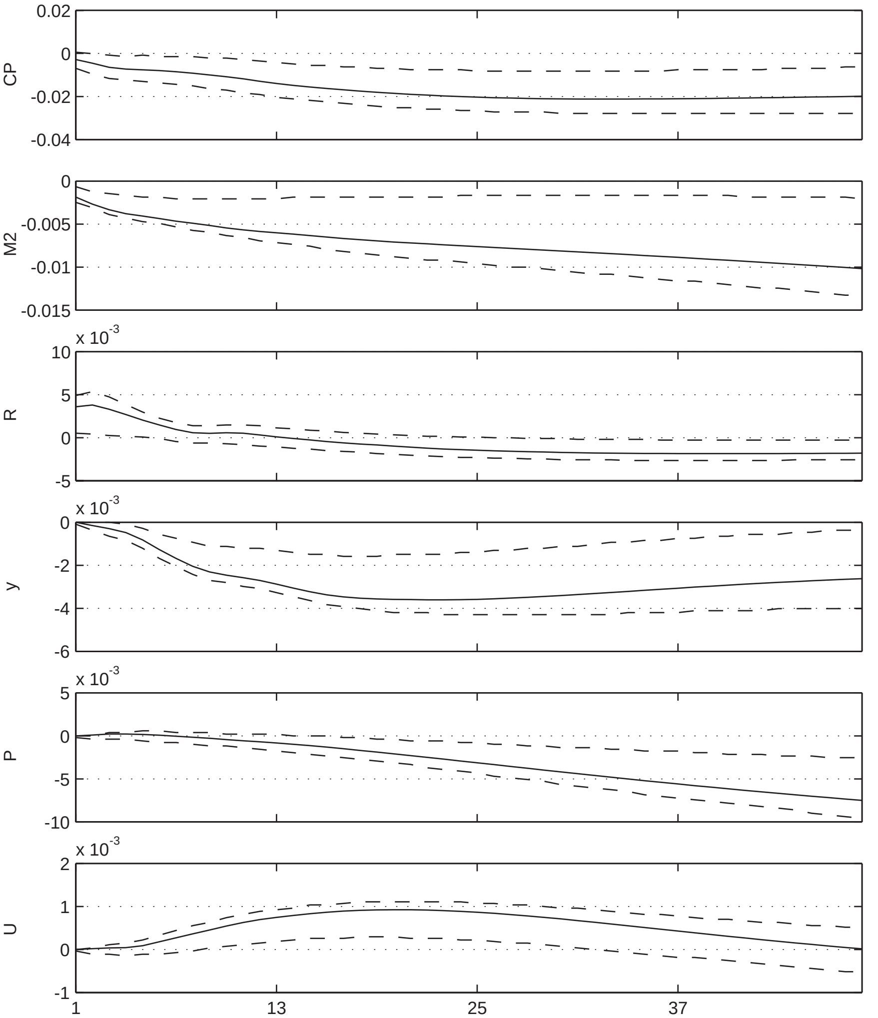

Fig. 2. Responses to an exogenous monetary policy contraction: monthly. Maximum likelihood estimates (solid) and 68% probability bands (dashed). Log levels or percentage points.

Elsewhere we evaluate in-sample and out-of-sample fits, check whether policy and nonpolicy disturbances are uncorrelated, and examine the stability of responses to exogenous changes in monetary policy (Leeper and Zha, 2002a). An exact small-sample test of the restricted model relative to the reduced-form model using a Bayes factor finds the data weakly favor the restricted model. The model's out-of-sample forecast performance compares favorably to various econometric models, including the Board of Governor's Greenbook forecasts.

A small-sample procedure finds the policy shock is uncorrelated with each of the nonpolicy shocks and with all five nonpolicy shocks jointly. With the disturbances uncorrelated, it is reasonable to perform policy evaluation by conditioning on a path of exogenous policy interventions while drawing nonpolicy disturbances independently.

Finally, the dynamic effects of policy are stable across time. Leeper and Zha (2002a) report responses to a policy shock that are computed for systems estimated over four sub-periods: 1959:1–1998:9 (entire sample); 1959:1-1979:9 (pre-Volcker); 1959:1–1982:12 (including nonborrowed reserves targeting); 1959:1–1998:9 with 1979:10–1982:12 eliminated (excluding nonborrowed reserves targeting). The qualitative responses to exogenous policy interventions are robust across sub-periods.

Some implications flow from finding stable parameters. First, over the post-1959 period either monetary policy has resided in a single regime or the various regimes have been too close to be detected statistically. Either implication is consistent with simulation results for parameters calibrated to U.S. data (Section 3.3.3). Second, there is no evidence that agents' beliefs about regime have changed in quantitatively important ways.

5. Some practical analysis of U.S. monetary policy

This section addresses some questions that Federal Reserve officials may have asked during the 1990s. To do so we compute policy projections under several alternative scenarios tied to actual U.S. policy experience. In each case we check whether the direct effects from the identified VAR are good approximations to the total effects of the contemplated interventions.

Our check involves computing an empirical modesty statistic in the same way we computed η^* for the theory in Section 3.

5.1. Measuring policy interventions

VAR forecasts capture only the direct effects of interventions. When an intervention generates large direct effects relative to their standard error, we infer the intervention is not modest. The theory in Section 3 suggests that immodest interventions may produce important expectations–formation effects that undermine the VAR's forecast accuracy.

In the VAR the K-period forecast errors given information at time T arising from the intervention $I_T = \{\tilde{\varepsilon}_{P,T+1}, \ldots, \tilde{\varepsilon}_{P,T+K}\}$ are

$$\eta_P(T,K) = \sum_{s=0}^{K-1} C_s(:,i)\tilde{\varepsilon}_{P,T+K-s}, \tag{26}$$

where the monetary policy equation is the ith equation in the system. As in Section 3.3.1, $\eta_P(T,K)$ is transformed to the standard normal variable, $\eta_P^*(T,K)$.

5.2. An immodest intervention

It is commonplace for central banks to condition forecasts on an unchanged path of the policy instrument. The Bank of England (2000), for example, projects GDP growth and inflation over horizons exceeding two years under the assumption that the official interest rate is constant. As the state of the economy is forecasted to change, however, an unchanged path of the instrument always requires some pattern of intervention.

In September 1990, the federal funds rate stood at 8.20 percent. Suppose the Fed were to hold the rate fixed at 8.20 percent over the next 4 years. Doing this requires a sequence of positive $\tilde{\varepsilon}_P$'s that increase over the forecast horizon, and average 1 1/2 standard deviations in year 3 and over 2 in the last year. This intervention implies large direct effects:

Variable	$\eta^*_P(90:9,48)$
y	−7.09
P	−4.51
U	6.69

The intervention is not modest. This situation is similar to Specifications A and C in Section 3.3.2 where each month the intervention is small, but the overall intervention is very persistent.

5.3. Modest policy interventions can matter

The efficacy of our procedure rests on whether conditioning on modest interventions can inform routine policy decisions. It turns out that this class of interventions is rich: it can generate economically meaningful shifts in the distributions of forecasted macro variables and clarify for policymakers the tradeoffs among alternative policies. The analysis also demonstrates that a complete probability model, which quantifies the uncertainties that Brainard (1967) emphasizes, can answer complex joint probability questions about the tradeoffs policymakers face.

We conduct the analysis through the eyes of a policymaker who has information about the economy through September 1990. Minutes of the October 2, 1990 FOMC meeting reveal that the Fed predicted a mild downturn in economic activity followed by a rapid resumption of moderate growth. Political developments in the Middle East, however, generated concerns about future oil prices and created uncertainty about the outlook for inflation. Although the domestic policy directive that emerged from the meeting sought "to maintain the existing degree of pressure on reserve position," several FOMC members dissented. One member favored immediate easing and three members opposed the FOMC's perceived leaning in favor of easing.

In light of the dissension among FOMC members, we consider two scenarios for the menu of conditional projections to present to policymakers. The first scenario conditions forecasts on the actual path of the federal funds rate from October 1990 to January 1991. An alternative scenario considers tighter policy over those four months.

Fig. 3 reports the actual time series, the out-of-sample forecasts conditional on the actual path of the funds rate, and 68 percent probability bands for the forecasts. The actual funds rate was 8.11 percent in October, 7.81 percent in November, 7.31 percent in December, and 6.91 percent in January. With that path of the funds rate, there is substantial probability that inflation will rise above 5 1/2 or 6 percent through 1993, real growth will fall below 1 percent in 1991, and unemployment will rise to near 7 percent through 1993 (left side of figure). Based on the path of the funds rate, it may appear that the Fed was concerned primarily about recession. As it happened, inflation fell to 3 percent by 1992, a recession occurred from July 1990 to March 1991 (according to NBER dating), and unemployment hit 7 1/2 percent in 1992. Of course, as the model's forecasts confirm, policymakers were unaware in October that the recession began three months earlier.

The FOMC minutes report that some policymakers were concerned about higher inflation. For those members an advisor might prepare forecasts conditional on tighter policy. The forecasts assume an intervention that raises the funds rate by 50 basis points in October (to 8.70 percent) and an additional 25 basis points over the period from November 1990 to January 1991. Ex-ante it appears that tighter policy would reduce the likelihood of higher inflation, but at the cost of raising the probability of negative real growth in 1991 (right side of Fig. 3). The point forecasts of output growth and unemployment, conditional on tighter monetary policy, come very close to the actual paths of the variables in 1991 and 1992. In spite of the sharp decline in the projected funds rate in 1991–1993, the modest intervention exerts a contractionary influence.[13]

Debate during the October 1990 FOMC meeting centered on the tradeoffs associated with alternative policy choices. To answer the concerns over economic slowdown and higher inflation, Table 2 reports a variety of joint probabilities involving real GDP growth in 1991 or 1992 or 1993 and inflation in 1992 and 1993, conditional on two alternative policies. "Tighter" policy assumes the same counterfactual policy behavior as in the right side of Fig. 3, while "Actual R^f" adjusts policy to be consistent with the left side.

The probabilities put a sharp point on the tradeoffs Federal Reserve officials perceived they faced. In terms of the marginal probabilities, tighter policy makes it very likely that inflation will remain low in 1992 and 1993 (below 5 1/2 percent), but

[13] To hit the actual path of the funds rate the intervention is $I_T = \{0.5, 0.1, -0.7, -0.7, 0, \ldots, 0\}$ and for tightening it is $I_T = \{2.3, 1.7, 1.0, 0.9, 0, \ldots, 0\}$. The resulting modesty statistics are

Variable	$\eta^*_P(90:9,48)$ Actual R^f	$\eta^*_P(90:9,48)$ Tighter Policy
y	0.10	−0.69
P	0.19	−1.49
U	−0.02	0.04

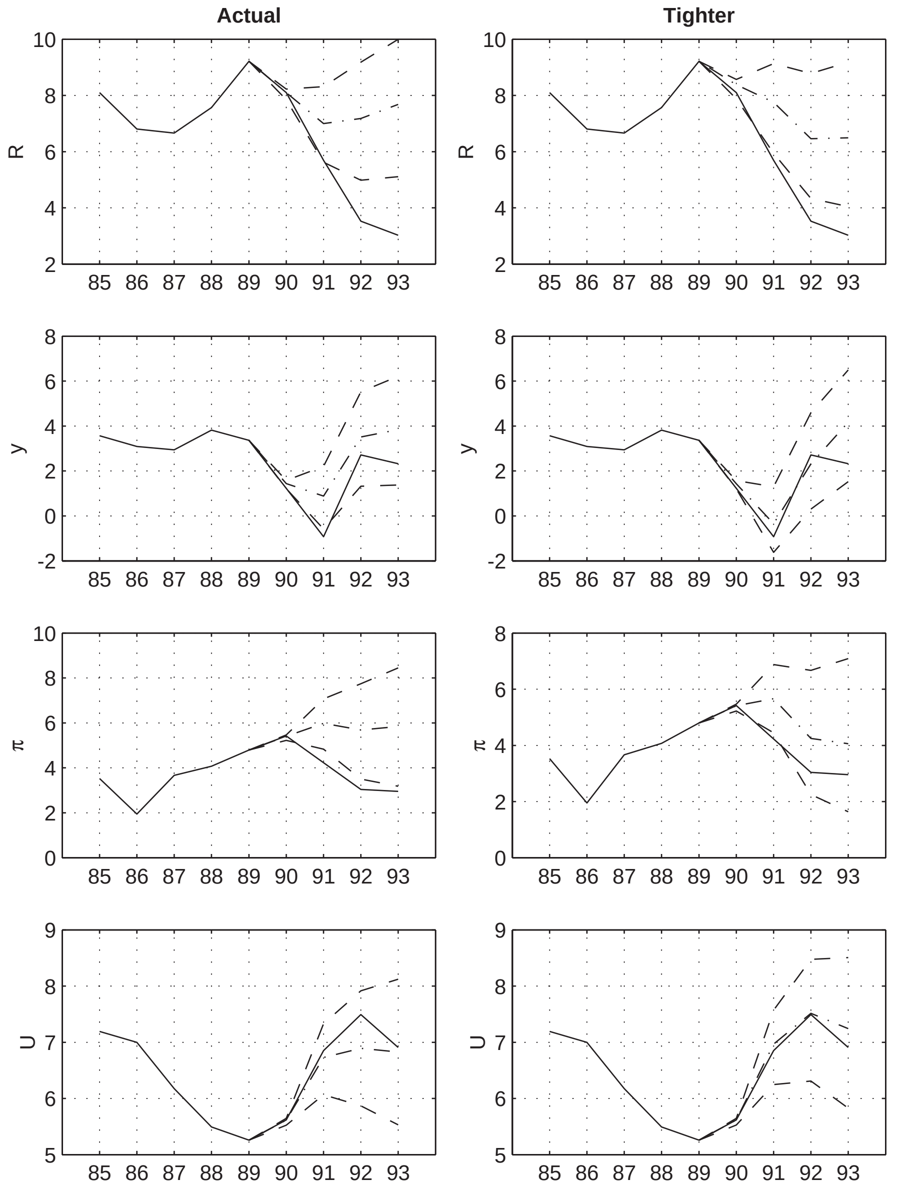

Fig. 3. Forecasts conditional on actual and tighter policy. Actual (solid) and out-of-sample forecast (dashed). First column: forecasts conditional on actual path of the federal funds rate from October 1990 to January 1991; second column: forecasts conditional on tighter policy (see text for details). Maximum likelihood estimates (dashed–dot) and 68% probability bands (dashed). Annual average growth rates or percentage points.

Table 2
Joint and marginal probabilities conditional on alternative policies

Outcomes based on out-of-sample forecasts from September 1990.
"Tighter" policy raises R^f to 8.70% in October and to 8.95% in November 1990–January 1991 and is produced by the sequence of exogenous actions $\tilde{\varepsilon}_P = (2.3, 1.7, 1.0, 0.9)$.
"Actual R^f" sets R^f at 8.11% in October, 7.81% in November, 7.31% in December, 6.91% in January 1991 and is produced by the sequence of exogenous actions $\tilde{\varepsilon}_P = (0.5, 0.1, -0.7, -0.7)$.

	Tighter	Actual R^f
P(*low* π *in* 1992)	0.67	0.47
P(*low* π *in* 1993)	0.66	0.46
P(*low* π *in* 1992 and 1993)	0.57	0.36
P(*recession in* 1991)	0.53	0.27
P(*recession in* 1992)	0.12	0.05
P(*recession in* 1993)	0.05	0.06
P(*recession* and *low* π)	0.33	0.11
P(*recession* and *high* π)	0.25	0.22
P(*no recession* and *low* π)	0.24	0.25
P(*no recession* and *high* π)	0.18	0.42

P(*recession*) is the probability of negative real GDP growth in 1991 or 1992 or 1993. P(*low* π) is the probability of inflation below 5 1/2 percent in 1992 and 1993. P(*recession* and *low* π) is the probability of negative real GDP growth in 1991 or 1992 or 1993 *and* inflation below 5 1/2 percent in 1992 and 1993.

it also produces a better than 50 percent chance of a recession in 1991 (negative real GDP growth for the year). More relevant to policymakers is the apparent tradeoff: tighter policy creates a one-third chance of a recession in 1991 or 1992 or 1993 and low inflation in 1992 and 1993.

For an intervention to generate the actual path of the funds rate, the Fed would have to tighten slightly in October and November and then ease in December and January. The column in the table labeled "Actual R^f" reports these results. This policy reduces by half the marginal probability of a recession in 1991 while lowering the marginal likelihoods of low inflation in 1992 and 1993. It also greatly reduces the joint probability of a recession in 1991 or 1992 or 1993 and low inflation in 1992 and 1993. Again the tradeoff is clear: the probability of no recession coupled with inflation over 5 1/2 percent now exceeds 40 percent, compared to 18 percent when policy is tighter.

5.4. Appraising and reappraising policy with modest interventions

Blinder's (1997) description of the appraisal/reappraisal process inherent in routine policymaking is echoed by Kohn (1995, p. 235) who observes that policymakers must "be flexible in revising forecasts and the policy stance in response to new information." This perspective helps to understand the Federal Reserve's "preemptive strike" against inflation in 1994. Our analysis shows that in February the tighter policy looked to be sufficient to offset higher inflation in 1996 and 1997; it was not sufficient by April, once three months of new information arrived. When we reappraise policy in April, a further tightening appears necessary

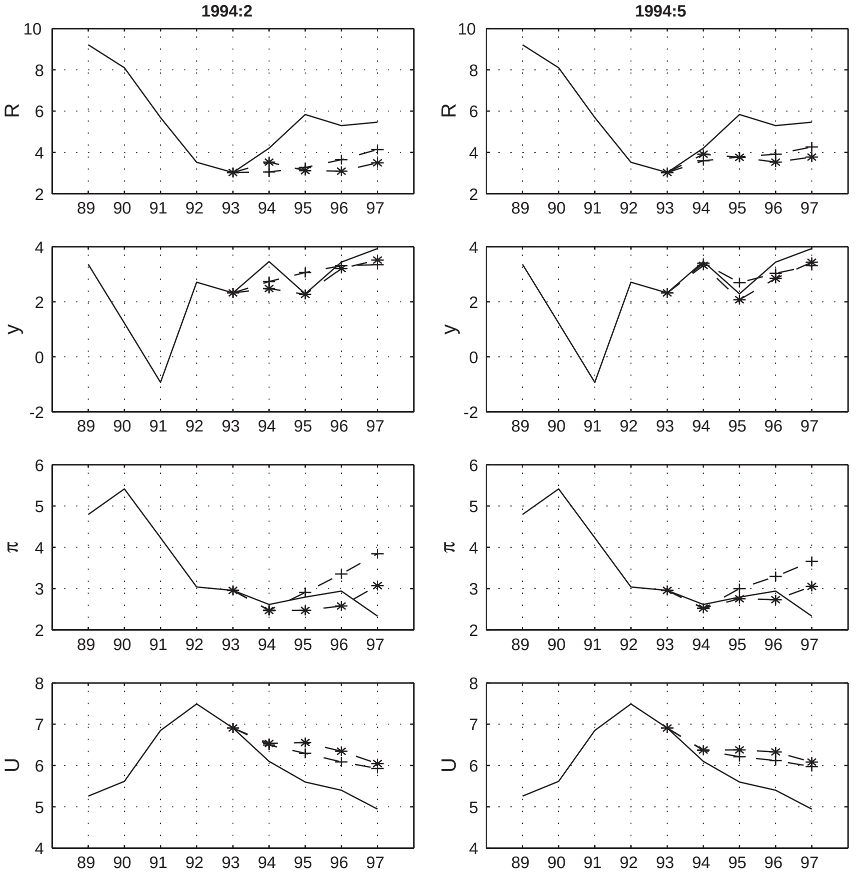

Fig. 4. Forecasts conditional on constant and actual funds rate in 1994. Actual (solid) and out-of-sample forecasts conditional on a constant funds rate (+) and on the actual path of the funds rate (∗) from February to May (first column) and from May to August (second column). In February, constant rate is 3.00%; actual path is 3.25%, 3.34%, 3.56%, 4.01%; in May, constant rate is 3.75%; actual path is 4.01%, 4.25%, 4.26%, 4.47%. Annual average growth rates or percentage points.

to preempt inflation. The analysis puts empirical flesh on Brainard (1967) argument for gradualism.

Fig. 4 displays actual data and out-of-sample forecasts made in January under two alternative policy scenarios for February through May. Given that the federal funds rate had been nearly constant at 3 percent over the previous year, a natural baseline maintains this constancy through May. The first intervention, which holds the funds rate constant for 4 months, is $I_T = \{-0.3, -0.2, -0.2, -0.2, 0, \ldots, 0\}$. That policy portends rising inflation over the next several years, exceeding 3 1/2 percent in 1996 and 1997; but policymakers do not seem to face an unpleasant tradeoff between inflation and real activity. Real GDP is expected to grow at least 3 percent annually from 1995 to 1997, while unemployment is projected to continue to decline.

Key policy questions at the time were when to raise the rate and how much to raise it. To address these questions we consider an alternative policy of moderate tightening that raises the funds rate along its actual path. Although the actual funds rate rose a full percentage point from January to May, that path requires a relatively small intervention, $I_T = \{0.5, 0, 0.5, 1.1, 0, \ldots, 0\}$.[14] In January even moderate tightening shifts the projected inflation path down without severely affecting real activity (left side of figure).

The Fed reappraises policy in April with three additional months of news about the economy. Now the baseline of a constant funds rate at 3.75 percent requires the intervention $I_T = \{0.1, -0.3, 0, \ldots, 0\}$, and leads policymakers to expect inflation will once again drift toward the 3 1/2 to 4 percent range in 1996 and 1997 (right side of Fig. 4). The outlooks for output and unemployment remain promising, so the Fed still does not face a difficult tradeoff. A somewhat stronger tightening move to match the actual path of the funds rate from May through August pushes the funds rate to 4.47 percent in August. The intervention $I_T = \{0.8, 0.4, 0.1, 0.8, 0, \ldots, 0\}$ achieves this funds rate path and shifts the mean forecast of inflation down below 3 percent through 1997 without risking recession.[15]

By reappraising their decisions in light of updated forecasts, policymakers move cautiously against inflation. So long as forecasts extend far enough into the future to capture monetary policy's lagged effects, the gradual approach can be successful. This analysis formalizes Blinder's (1997) description of policymaking. It also illustrates why uncertainty about future exogenous disturbances may lead policymakers to move cautiously, as Brainard (1967) instructs.[16]

[14] Both interventions are modest:

Variable	$\eta_P^*(94:1, 48)$ Constant R^f	$\eta_P^*(94:1, 48)$ Actual R^f
Y	0.12	−0.28
P	0.23	−0.53
U	−0.02	0.07

[15] Again the interventions are modest:

Variable	$\eta_P^*(94:4, 48)$ Constant R^f	$\eta_P^*(94:4, 48)$ Actual R^f
Y	0.03	−0.28
P	0.05	−0.54
U	−0.01	0.05

[16] Some readers might worry that if policymakers base their decisions on the procedure that we describe, agents' beliefs about regime might shift. Leeper and Zha (2002b) perform an ex-post check of how likely it was in January 1994 that expectations–formation effects were big even though the direct effects were small; an ex-post modesty statistic provides no warning signal.

6. Concluding remarks

This paper offers a framework for the analysis of modest policy interventions. The framework embeds rational expectations and respects the Lucas critique. In an environment where policy can switch regime and agents are continually learning about policy regime, the effects of a policy intervention consist of direct effects (which hold agents' beliefs about policy regime fixed) and expectations–formation effects (which reflect changes in expectations functions induced by changes in beliefs). When expectations–formation effects are important, agents' decision rules shift in precisely the way Lucas emphasizes. This behavior introduces a type of nonlinearity that can undermine the forecast accuracy of a linear econometric model. A modest policy intervention is a change in policy consistent with the historical variation in policy under the prevailing regime. Theoretical simulations show that an intervention can generate large direct effects without producing important expectations–formation effects.

The paper applies the theoretical framework to an identified VAR model of U.S. monetary policy. That model finds to be modest a rich class of interventions that the Federal Reserve routinely considers. The empirical model predicts that modest interventions can matter: they may shift the projected paths and probability distributions of macrovariables in economically meaningful ways. The empirical findings imply that accurately identified linear econometric models can reliably predict the impacts of modest policy interventions.

Our methodology has broad applicability. We have stressed its usefulness for the kinds of practical analyses central banks conduct. Closely related techniques are being used to study the behavior of the central bank of Sweden as it implements inflation targeting (Jansson and Vredin, 2001). Our method also yields insights about the plausibility of answers to counterfactual questions. Hamilton and Herrera (2002) apply the methodology to assess the plausibility of Bernanke et al. (1997) counterfactual analysis of the Federal Reserve's response to oil price shocks.

Nothing in the logic of our approach hinges on the use of an identified VAR as the forecasting tool. The distinction between direct effects and expectations–formation effects, the statistic we propose to check if an intervention is modest, and the pitfalls we enunciate all apply with equal force to any linear econometric model of policy.

Data appendix

The data, from 1959:1 to 1998:9, are collected from the Bureau of Economic Analysis, the Department of Commerce unless otherwise stated.

Federal Funds Rate: effective rate, monthly average. Source: Board of Governors of the Federal

Reserve System (BOG).

$M2$: $M2$ money stock, seasonally adjusted, billions of dollars. Source: BOG.

CPI: consumer price index for urban consumers (CPI-U), seasonally adjusted.

GDP: Real GDP, seasonally adjusted, billions of chain 1992 dollars. Monthly real GDP is interpolated using the procedure described in Leeper et al. (1996).

Unemployment: civilian unemployment rate (16 and over), seasonally adjusted. Source: Bureau of Labor Statistics.

Commodity prices: International Monetary Fund's index of world commodity prices. Source: *International Financial Statistics.*

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