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The Neo-Fisher Effect: Econometric Evidence from Empirical and Optimizing Models[†]

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This paper assesses the presence and importance of the neo-Fisher effect in postwar data. It formulates and estimates an empirical and a New Keynesian model driven by stationary and nonstationary monetary and real shocks. In accordance with conventional wisdom, temporary increases in the nominal interest rate are estimated to cause decreases in inflation and output. The main finding of the paper is that permanent monetary shocks that increase the nominal interest rate and inflation in the long run cause increases in interest rates, inflation, and output in the short run and explain about 45 percent of inflation changes. (*JEL* E12, E23, E31, E43, E52)

In the past two decades, a number of countries have been experiencing chronic below-target rates of inflation and near-zero nominal rates. According to the classic Fisher effect, nominal rates and inflation move together in the long run. This positive association is a robust empirical regularity. A less studied empirical question, however, is how a normalization of nominal interest rates (changes in the policy nominal interest rate that are expected to last for long periods of time) affects interest rates and inflation in the short run. This question is of interest because it can provide guidance on how monetary authorities can reflate their economies to levels consistent with their intended inflation targets. The present investigation addresses this question from an econometric perspective.

To this end, the paper develops a latent variable empirical model driven by transitory and permanent monetary and real shocks, then estimates it using Bayesian techniques on postwar data. Like DSGE models, the proposed empirical model allows for more structural shocks than time series, but with the advantage of requiring fewer structural restrictions.

In accordance with conventional wisdom, the estimated model predicts that a transitory increase in the nominal interest rate causes a fall in inflation, a contraction in real activity, and a rise in the real interest rate. The main result of the paper is that in response to a permanent monetary shock that increases the nominal interest rate and inflation in the long run, these two variables increase in the short run, reaching

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their higher long-run levels within a year. Furthermore, the adjustment to a permanent increase in the policy rate entails no output loss and is characterized by low real interest rates. Permanent monetary shocks are estimated to be the main drivers of inflation, explaining more than half of observed movements in the price level at business cycle frequency. These results represent the first econometric assessment of the presence and importance of the neo-Fisher effect in the data.

The paper then introduces nonstationary and stationary but persistent inflation target shocks into a standard optimizing New Keynesian model in which the central bank follows a Taylor-type interest rate feedback rule. In the model, the permanent and stationary inflation target shocks compete with standard transitory monetary shocks, permanent and transitory productivity shocks, a preference shock, and a labor supply shock. The goal of this analysis is not theoretical in nature. A number of papers, many of which are cited below, have demonstrated that in the New Keynesian model, sufficiently persistent movements in the inflation target are accommodated through rising interest rates and inflation in the short run. Instead, the objective of the analysis is to estimate the importance of shocks that give rise to this type of dynamic. The estimated New Keynesian model predicts that 50 percent of the variance of inflation changes is explained by monetary shocks that produce neo-Fisherian dynamics.

Taken together, the predictions of the estimated empirical and optimizing models suggest that there is a sizable neo-Fisher effect in the data. From a policy perspective, this result provides econometric support to the prediction that in a country facing below-target inflation and a near-zero nominal interest rate, a permanent increase in the rate of inflation is implemented via a credible normalization of the policy rate.

A byproduct of the econometric analysis conducted in this paper is the finding that distinguishing temporary and permanent monetary disturbances provides a resolution of the well-known price puzzle, according to which a transitory increase in the nominal interest rate is estimated to cause a short-run increase in inflation.

The neo-Fisherian approach pursued in the present investigation, according to which the inflation target has an exogenous nonstationary component, is clearly not the only possible interpretation of the joint long-run behavior of interest rates and inflation. At least two alternative views are a priori equally plausible. One maintains that the permanent component of inflation and nominal rates is not exogenous but is driven by other factors, such as public debt and the stream of current and future expected primary fiscal deficits, that ultimately determine prices and the monetary stance. Under this view, the steady increase of inflation and interest rates that started in the early 1960s and culminated with the Volcker disinflation as well as the subsequent gradual fall in these two variables over the Alan Greenspan and Ben Bernanke eras would be the result of not exogenous adjustments in the permanent component of the inflation target but rather the consequence of expansionary and contractionary (fiscal) policies dominating, respectively, the pre- and post-Volcker subsamples (Sims 2011). A second alternative view is a familiar one among monetary economists (e.g., Sargent 1999). It holds that the rise in inflation in the 1970s was the result of systematic overexpansionary monetary policy that eventually lost control of inflation and was then forced to raise policy rates persistently. According to this view, at some point during Volcker's tenure, policy reacted vigorously by aggressively increasing policy rates, which in turn generated a temporary recession and a declining path of inflation and, subsequently, of interest rates. These two alternative interpretations of the observed comovement of interest rates and inflation and the one provided in this paper are not necessarily mutually exclusive. For example, as shown in Section IIIA, the model proposed in this paper interprets the Volcker disinflation as a combination of a temporary increase in the nominal interest rate and a simultaneous gradual descent in its permanent component.

This paper is related to a number of theoretical and empirical contributions on the effects of interest rate policy on inflation and aggregate activity. On the empirical front, it is related to papers that estimate the short-run effects of permanent monetary shocks. Azevedo, Ritto, and Teles (2019), using a vector error correction model approach, confirm the results of this paper and add novel additional evidence for the United Kingdom, France, Germany, and the eurozone. Aruoba and Schorfheide (2011) estimate a model that combines New Keynesian and monetary search frictions. The permanent component of inflation predicted by their model is in line with the one estimated in this paper and gives rise to positive short-run comovement of inflation and interest rates. They show that the predictions of their model are consistent with those of an estimated vector autoregression (VAR) system. Gao, Kulish, and Nicolini (2020) estimate time-varying permanent components of inflation and the nominal rate and find that they comove closely in the short run. De Michelis and Iacoviello (2016) estimate an structural VAR (SVAR) model with permanent monetary shocks to evaluate the Japanese experience with Abenomics. They also study the effect of monetary shocks in the context of a calibrated New Keynesian model. The present paper departs from their work in two important dimensions. First, their SVAR model does not include the short-run policy rate. The inclusion of this variable is key in the present paper because the short-run comovement of the policy rate with inflation is at the core of the neo-Fisher effect. Second, their theoretical model is not estimated and does not include permanent monetary shocks. By contrast, this paper allows permanent and transitory monetary shocks to compete with each other and with other shocks in the econometric estimation, and as pointed out above, it finds that permanent monetary shocks are important drivers of movements in inflation. Fève, Matheron, and Sahuc (2010) estimate SVAR and dynamic optimizing models with nonstationary inflation target shocks to study the role of gradualism in disinflation policy. They show, by means of counterfactual experiments, that had the European monetary authority been less gradual in lowering its inflation target during the late 2000s, the eurozone would have suffered a milder slowdown in economic growth. The present paper focuses instead on how the short-run comovement of inflation and the policy rate triggered by a monetary disturbance change depending on whether the impulse is permanent or transitory in nature. King and Watson (2012) find that in estimated New Keynesian models, postwar US inflation is explained mostly by variations in nonstandard shocks, such as random variations in markups. The present paper shows that once one allows for permanent monetary shocks, almost half of the variance of inflation changes is explained by monetary disturbances. Sims and Zha (2006) estimate a regime-switching model for US monetary policy and find that during the postwar period, there were three policy regime switches, but they were too small to explain the observed increase in inflation of the 1970s or the later disinflation that started with the Volcker chairmanship. The empirical and optimizing models estimated in the present paper attribute much of the movement in inflation in these two episodes to the permanent nominal shock. Cogley and Sargent (2005) use an autoregressive framework to produce estimates of long-run inflationary expectations. The predictions of the two models estimated in the present paper are consistent with their estimates of long-run inflation expectations.

This paper is also related to a body of work that incorporates inflation target shocks in the New Keynesian model. In this regard, the contribution of the present paper is to allow for a permanent component in this source of inflation dynamics. Garín, Lester, and Sims (2018) show that the New Keynesian model delivers neo-Fisherian effects in response to increases in the inflation target, provided the latter are sufficiently persistent. They also show that the neo-Fisher effect weakens as firms become more backward looking in their pricing behavior. The present investigation is complementary to this work by providing econometric estimates of both the persistence of the inflation target shock and the backward-looking component in the price-setting mechanism. It shows that the estimated parameters give rise to neo-Fisherian dynamics in response to innovations in the stationary component of the inflation target. It also finds that this shock explains a sizable fraction of the variance of changes in the inflation rate. Ireland (2007) estimates a New Keynesian model with a time-varying inflation target and shows that, possibly as a consequence of the Federal Reserve's attempt to accommodate supply-side shocks, the inflation target increased significantly during the 1960s and 1970s and fell sharply in the early 2000s. Using a similar framework as Ireland's, Milani (2020) shows that movements in the inflation target become less pronounced if one assumes that agents must learn about the level of the inflation target.

This paper is also related to recent theoretical developments on the neo-Fisher effect. Schmitt-Grohé and Uribe (2010, 2014) show that the neo-Fisher effect obtains in the context of standard dynamic optimizing models with flexible prices. Specifically, they show that a credible increase in the nominal interest rate that is expected to be sustained for a prolonged period of time gives rise to an immediate increase in inflationary expectations. Schmitt-Grohé and Uribe (2010, 2014) show that this result also obtains in models with nominal rigidity. Cochrane (2017) shows that if the monetary policy regime is passive, a temporary increase in the nominal interest rate can cause an increase in the short-run rate of inflation. This notion of the neo-Fisher effect is different from the one studied in the present paper, which associates the neo-Fisher effect with the short-run response of inflation to monetary shocks that move inflation and interest rates in the long run. Williamson (2018) considers a model with flexible-price and sticky-price goods and shows that movements in the interest rate generate movements in expected flexible-price inflation of equal size. Cochrane (2014) and Williamson (2016) provide nontechnical expositions of the neo-Fisher effect. Finally, Lukmanova and Rabitsch (2021) extend the analysis in the present paper by incorporating imperfect information along the lines of Erceg and Levin (2003). They find that in response to a persistent increase in the inflation target, the neo-Fisher effect takes place with some delay.

The remainder of the paper proceeds as follows. Section I presents evidence consistent with the long-run validity of the Fisher effect. Section II presents the proposed empirical model and discusses the identification and estimation strategies. Section III presents the estimated short-run effects of permanent monetary shocks on inflation, the interest rate, and output. It also reports the importance of these shocks in explaining changes in the rate of inflation. Section IV presents the New Keynesian model and the estimated effects of permanent and stationary monetary shocks. Section V closes the paper with a discussion of actual monetary policy in the ongoing low-inflation era from the perspective of the two estimated models. Data and replication code are available online at the journal's official repository and on the author's website.

I. Evidence on the Fisher Effect

What is the effect of an increase in the nominal interest rate on inflation? One can argue that the answer to this question depends on (i) whether the increase in the interest rate is expected to be permanent or transitory and (ii) whether the horizon of interest is the short run or the long run. Thus, the question that opens this section represents, in fact, four questions. Table 1 summarizes the state of the monetary literature in the quest to answer them.

A large body of empirical and theoretical studies argues that a transitory positive disturbance in the nominal interest rate causes a transitory increase in the real interest rate, which in turn depresses aggregate demand and inflation, entry $(1,2)$ in the table (see, for example, Christiano, Eichenbaum, and Evans 2005—henceforth, CEE—2005, especially Figure 1). Similarly, a property of virtually all modern models studied in monetary economics is that a transitory increase in the nominal interest rate has no effect on inflation in the long run, entry $(1,1)$. By contrast, if the increase in the nominal interest rate is permanent, then sooner or later, inflation will have to increase by roughly the same magnitude if the real interest rate, given by the difference between the nominal rate and expected inflation, is not determined by nominal factors in the long run, entry $(2,1)$ in the table. This long-run relationship between nominal rates and inflation is known as the Fisher effect. Until recently, there was no answer to the question of how a monetary shock that increases interest rates and inflation in the long run should affect these variables in the short run, entry $(2,2)$ in the table. The relatively novel neo-Fisher effect says that a permanent increase in the nominal interest rate causes an increase in inflation in not only the long run but also the short run, so that entry $(2,2)$ in the table should be a plus sign. Thus far, there exists no formal empirical analysis of this effect. The focus of the present investigation is to fill this gap by ascertaining whether the neo-Fisher effect is present in the data.

Before plunging into an econometric analysis of the neo-Fisher effect, I wish to briefly present evidence consistent with the Fisher effect. The rationale for doing so is that my empirical analysis of the neo-Fisher effect assumes the empirical validity of the Fisher effect, interpreted as a long-run positive relationship between the nominal interest rate and inflation. The left panel of Figure 1 displays time averages of inflation and nominal interest rates across 99 countries. Each dot in the graph

Table 1—Effect of an Increase in the Nominal Interest Rate on Inflation

	Long run effect	**Short** run effect
Transitory shock	0	$\downarrow$
Permanent shock	$\uparrow$	$\boxed{\uparrow ?}$

Notes: Entry (2, 1): The Fisher effect. Entry (2, 2): The Neo-Fisher effect.

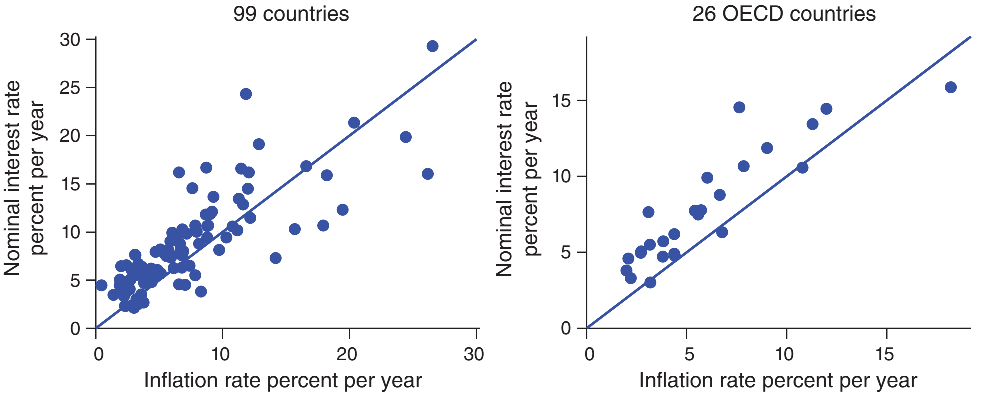

Figure 1. Average Inflation and Nominal Interest Rates: Cross-Country Evidence

Notes: Each dot represents one country. For each country, averages are taken over the longest available uninterrupted sample. The average sample covers the period 1989 to 2012. The solid line is the 45-degree line.

Sources: World Development Indicators (WDI), available at data.worldbank.org/indicator. Inflation is the CPI inflation rate (code FP.CPI.TOTL.ZG). The nominal interest rate is the treasury bill rate. The WDI database provides this time series not directly, but as the difference between the lending interest rate (code FR.INR.LEND) and the risk premium on lending (lending rate minus treasury bill rate, code FR.INR.RISK). Countries for which one or more of these series were missing as well as outliers, defined as countries with average inflation or interest rate above 50 percent, were dropped from the sample.

corresponds to one country. The typical sample covers the period 1989 to 2012. The scatter plot is consistent with the Fisher effect in the sense that increases in the nominal interest rate are associated with increases in the rate of inflation. This is also the case for the subsample of Organisation for Economic Co-operation and Development countries (right panel), which are on average half as inflationary as the group of nonmember countries. Figure 2 presents empirical evidence consistent with the Fisher effect from the time perspective. It plots inflation and the nominal interest rate in the United States over the period 1954:IV to 2018:II. In spite of the fact that the data have a quarterly frequency, it is possible to discern a positive long-run association between inflation and the nominal rate. This relation becomes even more apparent if one removes the cyclical component of both series as in Gao, Kulish, and Nicolini (2020), who separate trend and cycle using the Hodrick–Prescott filter. The high inflations of the 1970s and 1980s coincided with high levels of the interest rate. Symmetrically, the relatively low rates of inflation observed since the early 1990s have been accompanied by low nominal rates.

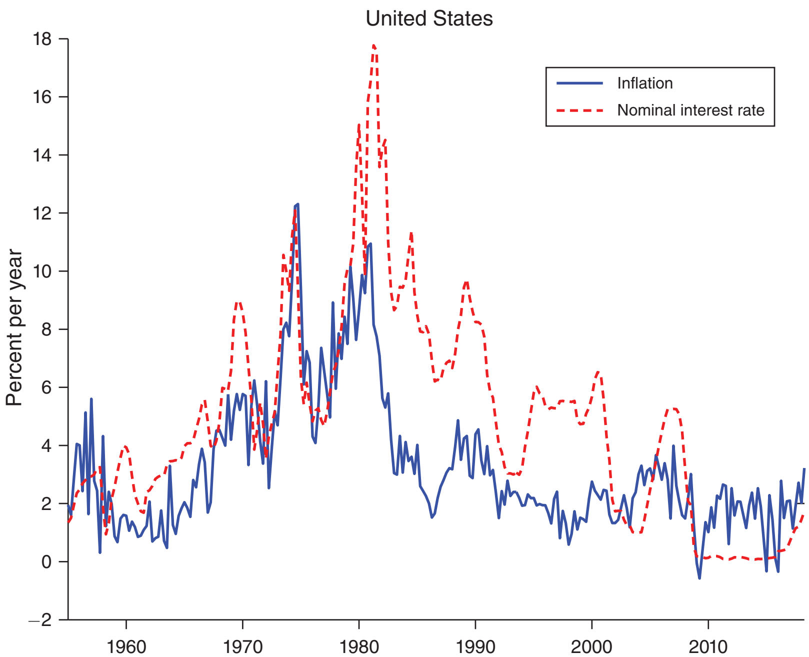

Figure 2. Inflation and the Nominal Interest Rate in the United States

Note: Quarterly frequency.

Source: See Section IIC

The Fisher effect, however, does not provide a prediction of when inflation should be expected to catch up with a permanent increase in the nominal interest rate. It only states that it must eventually do so. A natural question, therefore, is how quickly does inflation adjust to a permanent increase in the nominal interest rate? The remainder of this paper is devoted to addressing this question.

II. The Empirical Model

The empirical model is a system of latent variables in the spirit of DSGE models, but with fewer cross-coefficient restrictions. It allows, for example, for more identified shocks than observable time series, thereby allowing for more flexibility than SVAR systems. The model aims to capture the dynamics of three macroeconomic indicators—namely, the logarithm of real output per capita, denoted y_t; the inflation rate, denoted π_t and expressed in percent per year; and the nominal interest rate, denoted i_t and also expressed in percent per year. Section IIIC extends the model to include the ten-year spread. I assume that y_t, π_t, and i_t are driven by four exogenous shocks: a nonstationary (or permanent) monetary shock, denoted X_t^m; a stationary (or transitory) monetary shock, denoted z_t^m; a nonstationary nonmonetary shock, denoted X_t; and a stationary nonmonetary shock, denoted z_t. The focus of my analysis is the short-run effects of innovations in X_t^m and z_t^m. The shocks X_t and z_t are meant to capture the nonstationary and stationary components of combinations of nonmonetary disturbances of different natures, such as technology shocks, preference shocks, or markup shocks, which my analysis is not intended to individually identify.

I assume that output is cointegrated with X_t and that inflation and the nominal interest rate are cointegrated with X_t^m. I then define the following vector of stationary variables:

$$\begin{bmatrix} \hat{y}_t \\ \hat{\pi}_t \\ \hat{i}_t \end{bmatrix} \equiv \begin{bmatrix} y_t - X_t \\ \pi_t - X_t^m \\ i_t - X_t^m \end{bmatrix}.$$

The variable $\hat{y}_t$ represents detrended output, and $\hat{\pi}_t$ and $\hat{i}_t$ represent the cyclical components of inflation and the nominal interest rate. Because inflation and the nominal interest rate share a common nonstationary component, they are cointegrated. Here, the cointegrating vector is $[1\ -1]$. Section IIIC relaxes this assumption to allow for nonstationarity in the real interest rate.

The law of motion of the vector $[\hat{y}_t\ \hat{\pi}_t\ \hat{i}_t]'$ is assumed to take the autoregressive form[1]

$$\begin{bmatrix} \hat{y}_t \\ \hat{\pi}_t \\ \hat{i}_t \end{bmatrix} = \sum_{i=1}^{4} B_i \begin{bmatrix} \hat{y}_{t-i} \\ \hat{\pi}_{t-i} \\ \hat{i}_{t-i} \end{bmatrix} + C \begin{bmatrix} \Delta X_t^m \\ z_t^m \\ \Delta X_t \\ z_t \end{bmatrix}, \tag{1}$$

where $\Delta X_t^m \equiv X_t^m - X_{t-1}^m$, $\Delta X_t \equiv X_t - X_{t-1}$, and B_i and C are matrices of coefficients to be estimated. The driving forces are assumed to follow univariate AR(1) laws of motion of the form

$$\begin{bmatrix} \Delta X_{t+1}^m \\ z_{t+1}^m \\ \Delta X_{t+1} \\ z_{t+1} \end{bmatrix} = \rho \begin{bmatrix} \Delta X_t^m \\ z_t^m \\ \Delta X_t \\ z_t \end{bmatrix} + \psi \begin{bmatrix} \epsilon_{t+1}^1 \\ \epsilon_{t+1}^2 \\ \epsilon_{t+1}^3 \\ \epsilon_{t+1}^4 \end{bmatrix}, \tag{2}$$

where ρ and ψ are diagonal matrices of coefficients to be estimated and ϵ_t^i are i.i.d. disturbances distributed $N(0,1)$.

A. *Identification*

Thus far, I have introduced three identification assumptions—namely, that output is cointegrated with X_t and that inflation and the interest rate are cointegrated with X_t^m. In addition, to identify the transitory monetary shock, z_t^m, I use two alternative strategies: The baseline strategy is to impose sign restrictions on the impact effect of these disturbances on endogenous variables. Specifically, I assume that

$$C_{12}, C_{22} \leq 0,$$

[1] The presentation of the model omits intercepts. A detailed exposition is in online Appendix A.

where C_{ij} denotes the (i,j) element of C. These two conditions restrict transitory exogenous increases in the interest rate to have nonpositive impact effects on output and inflation. The alternative identification strategy, pursued in Section IIIC, is to assume that stationary monetary shocks have no impact effect on output and inflation,

$$C_{12} = C_{22} = 0.$$

Both schemes yield similar results. As explained in Section IIE, additional identification restrictions aimed at distinguishing z_t^m from z_t are imposed via restrictions on the prior distributions of the elements of C and ρ. Finally, without loss of generality, I introduce the normalizations $C_{32} = C_{14} = 1$.

B. *Observables*

All variables in the system (1)–(2) are unobservable. To estimate the parameters of the matrices defining this system, I use observable variables for which the model has precise predictions. Specifically, I use observations of output growth, Δy_t; the change in the nominal interest rate, Δi_t; and the interest rate–inflation differential,

$$r_t \equiv i_t - \pi_t.$$

These three variables are stationary by the maintained long-run identification assumptions. The following equations link the observables to variables included in the unobservable system (1)–(2):

$$\begin{aligned} \Delta y_t &= \hat{y}_t - \hat{y}_{t-1} + \Delta X_t, \\ r_t &= \hat{i}_t - \hat{\pi}_t, \\ \Delta i_t &= \hat{i}_t - \hat{i}_{t-1} + \Delta X_t^m. \end{aligned} \tag{3}$$

As in much of the literature on estimation of dynamic macroeconomic models using Bayesian techniques, I assume that Δy_t, r_t, and Δi_t are observed with measurement error. Formally, letting o_t be the vector of variables observed in quarter t, I assume that

$$o_t = \begin{bmatrix} \Delta y_t \\ r_t \\ \Delta i_t \end{bmatrix} + \mu_t, \tag{4}$$

where μ_t is a vector of measurement errors distributed i.i.d. $N(\varnothing, R)$ and R is a diagonal variance-covariance matrix. These shocks play a role similar to that of regression residuals in classic estimation. As explained in more detail below, measurement errors are restricted to explain no more than 10 percent of the variance of the observables. The main results of the paper are robust to doing away with measurement errors.

C. *The Data*

I estimate the empirical model on quarterly US data spanning the period 1954:III to 2018:II. The proxy for y_t is the logarithm of real GDP seasonally adjusted in chained dollars of 2012 (US Bureau of Economic Analysis 1947–2018b) minus the logarithm of the civilian noninstitutional population 16 years old or older (US Bureau of Labor Statistics 2018). The proxy for π_t is the growth rate of the implicit GDP deflator expressed in percent per year. In turn, the implicit GDP deflator is constructed as the ratio of GDP in current dollars (US Bureau of Economic Analysis 1947–2018a) and real GDP, both seasonally adjusted. The proxy for i_t is the monthly federal funds effective rate (Board of Governors of the Federal Reserve System 1957–2018), converted to quarterly frequency by averaging and expressed in percent per year.

D. *Estimation*

The model is estimated using Bayesian techniques. To compute the likelihood function, it is convenient to use the state-space representation of the model. Define the vector of endogenous variables $\hat{Y}_t \equiv [\hat{y}_t \ \hat{\pi}_t \ \hat{i}_t]'$ and the vector of driving forces $u_t \equiv [\Delta X_t^m \ z_t^m \ \Delta X_t \ z_t]'$. Then, the state of the system is given by

$$\xi_t \equiv \begin{bmatrix} \hat{Y}_t \\ \hat{Y}_{t-1} \\ \vdots \\ \hat{Y}_{t-3} \\ u_t \end{bmatrix},$$

and the system composed of equations (1)–(4) can be written as follows:

$$\xi_{t+1} = F\xi_t + P\epsilon_{t+1}$$

$$o_t = H'\xi_t + \mu_t,$$

where the matrices F, P, and H are known functions of B_i, $i = 1,\ldots 4$, C, ρ, and ψ and are presented in online Appendix A. This representation allows for the use of the Kalman filter to evaluate the likelihood function, which facilitates estimation.

E. *Priors*

Table 2 displays the prior distributions of the estimated coefficients. The prior distributions of all elements of B_i, for $i = 1, \ldots, 4$, are assumed to be normal. In

Table 2—Prior Distributions

Parameter	Distribution	Mean	Standard deviation
Diagonal elements of B_1	Normal	0.95	0.5
All other elements of B_i, $i = 1, 2, 3, 4$	Normal	0	0.25
C_{21}, C_{31}	Normal	-1	1
$-C_{12}, -C_{22}$	Gamma	1	1
All other estimated elements of C	Normal	0	1
ψ_{ii}, $i = 1, 2, 3, 4$	Gamma	1	1
ρ_{ii}, $i = 1, 2, 3$	Beta	0.3	0.2
ρ_{44}	Beta	0.7	0.2
R_{ii}, $i = 1, 2, 3$	$\text{Uniform}\left[0, \frac{\text{var}(o_t)}{10}\right]$	$\frac{\text{var}(o_t)}{10 \times 2}$	$\frac{\text{var}(o_t)}{10 \times \sqrt{12}}$

the spirit of the Minnesota prior (MP), I assume a prior parameterization in which at the mean of the prior distribution, the elements of $\hat{Y}_t$ follow univariate autoregressive processes. So when evaluated at their prior means, only the diagonal elements of B_1 take nonzero values, and all other elements of B_i for $i = 1, \ldots, 4$ are nil. Because the system (1)–(2) is cast in terms of stationary variables, I deviate from the random walk assumption of the MP and instead impose an autoregressive coefficient of 0.95 in all equations, so that all elements along the main diagonal of B_1 take a prior mean of 0.95. I assign a prior standard deviation of 0.5 to the diagonal elements of B_1, which implies a coefficient of variation close to one half (0.5/0.95). As in the MP, I impose lower prior standard deviations on all other elements of the matrices B_i for $i = 1, \ldots, 4$ and set them to 0.25.

The coefficient C_{21} takes a normal prior distribution with mean -1 and standard deviation 1. This implies a prior belief that the impact effect of a permanent interest rate shock on inflation, given by $1 + C_{21}$, can be positive or negative with equal probability. I make the same assumption about the impact effect of permanent monetary shocks on the nominal interest rate itself, thus assigning to C_{31} a normal prior distribution with mean -1 and standard deviation 1. Under the baseline identification scheme for the transitory monetary shock z_t^m, $-C_{12}$, and $-C_{22}$ are restricted to be nonnegative. I assume that they have gamma prior distributions with mean and standard deviations equal to 1. All other parameters of the matrix C, except for C_{32} and C_{14} (which are normalized to unity), are assigned a normal prior distribution with mean 0 and standard deviation 1.[2] The parameters

[2]One might wonder whether a rationale like the one I used to set the prior mean of C_{21} could apply to C_{13}, the parameter governing the impact output effect of a nonstationary nonmonetary shock, X_t, which is given by $1 + C_{13}$. To see why a prior mean of 0 for C_{13} might be more reasonable, consider the effect of an innovation in the permanent component of TFP, which is perhaps the most common example of a nonstationary nonmonetary shock in business cycle analysis. Specifically, consider a model with the Cobb-Douglas production function $y_t = X_t + z_t + \alpha k_t + (1 - \alpha) h_t$ expressed in logarithms. Consider first a situation in which capital and labor, denoted k_t and h_t, do not respond contemporaneously to changes in X_t. In this case, the contemporaneous effect of a unit increase in X_t on output is unity, which implies that a prior mean of 1 for $1 + C_{13}$, or, equivalently, a prior mean of 0 for C_{13} is the most appropriate. Now consider the impact effect of changes in X_t on k_t and h_t. It is reasonable to assume that the stock of capital, k_t, is fixed in the short run. The response of h_t depends on substitution and wealth effects. The former tends to cause an increase in employment, and the latter a reduction. Which effect will prevail is

ψ_{ii}, for $i = 1,\ldots,4$, representing the standard deviations of the four exogenous innovations in the AR(1) process (2), are all assigned gamma prior distributions with mean and standard deviation equal to 1. I adopt beta prior distributions for the serial correlations of the driving processes, ρ_{ii}, $i = 1,\ldots,4$. I assume relatively small means of 0.3 for the prior serial correlations of the 2 monetary shocks and the nonmonetary nonstationary shock and assume a relatively high mean of 0.7 for the stationary nonmonetary shock. The small prior mean serial correlations for the monetary shocks reflect the usual assumption in the related literature of serially uncorrelated monetary shocks. The relatively small prior mean serial correlation for the nonstationary nonmonetary shock reflects the fact that the growth rate of the stochastic trend of output is typically estimated to have a small serial correlation. Similarly, the relatively high prior mean of the serial correlation of the stationary nonmonetary shock reflects the fact that typically these shocks (e.g., the stationary component of TFP) are estimated to be persistent. The prior distributions of all serial correlations are assumed to have a standard deviation of 0.2.

The restrictions imposed on the prior distributions of the elements of the matrices C and ρ play a role in the identification of z_t^m and z_t in not only the statistical sense but also, and more importantly, the economic sense. Interestingly, the assumed identification scheme allows for the possibility of a second stationary monetary shock, like in the New Keynesian DSGE model of Section IV. This would be the case if the estimate of C_{24} is positive and that of C_{34} is negative (recall that C_{14} is normalized to 1). In this case, the prior restrictions on C and ρ guarantee that the two stationary monetary shocks would be distinct. For example, the shock z_t will tend to be more persistent than z_t^m (recall that their mean prior serial correlations are 0.7 and 0.3, respectively) and would have the interpretation of a stationary shock to the inflation target, as in much of the literature on trend inflation. As it turns out, the actual estimate of z_t is not of this type. I will continue to refer to z_t as the nonmonetary stationary shock because ex ante, only z_t^m is guaranteed to be a stationary monetary shock as defined here.

The variances of all measurement errors are assumed to have a uniform prior distribution with lower bound 0 and an upper bound of 10 percent of the sample variance of the corresponding observable indicator.

Finally, to draw from the posterior distribution of the estimated parameters, I apply the Metropolis-Hastings sampler to construct a Monte Carlo Markov chain (MCMC) of 1 million draws after burning the initial 100,000 draws. Posterior means and error bands around the impulse responses shown in later sections are constructed from a random subsample of the MCMC chain of length 100,000 with replacement.

To check for the identifiability of the estimated parameters of the model, I apply the test proposed by Iskrev (2010). This procedure consists in calculating the derivative of the predicted autocovariogram of the observables with respect to the vector of estimated parameters. Identifiability obtains if the matrix of derivatives

not clear, giving credence to a prior of 0 for C_{13}. One could further think about the role of variable input utilization. An increase in X_t is likely to cause an increase in utilization, further favoring a prior mean of 0 over one of -1 for C_{13}.

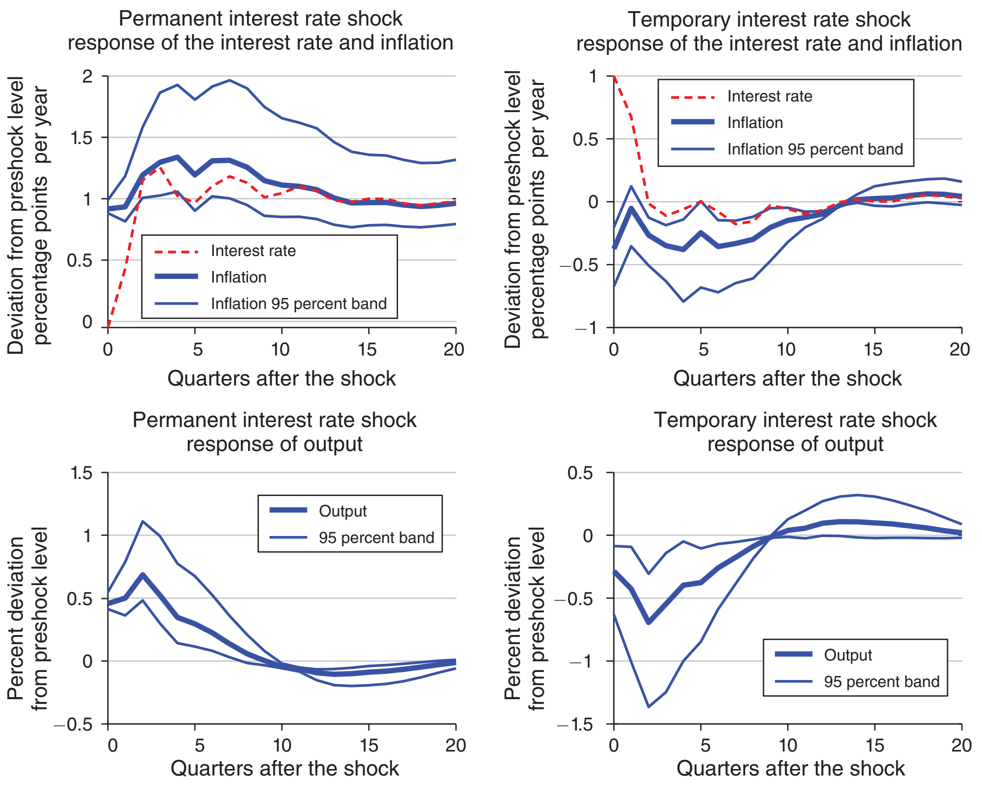

Figure 3. Impulse Responses to Permanent and Temporary Interest Rate Shocks: Empirical Model

Notes: Impulse responses are posterior mean estimates. Asymmetric error bands are computed using the Sims-Zha (1999) method.

has rank equal to the length of the vector of estimated parameters. Evaluating the parameters of the model at their posterior mean, I find that the rank condition is satisfied. This means that in a neighborhood of the posterior mean, the predicted covariogram is uniquely determined by the value of the vector of estimated parameters.

III. Effects of Permanent and Transitory Monetary Shocks

Figure 3 displays mean posterior estimates of the responses of inflation, output, and the nominal interest rate to a permanent monetary shock (an increase in X_t^m) and a temporary interest rate shock (an increase in z_t^m). The size of the permanent monetary shock is set to ensure that on average, it leads to a 1 percent increase in the nominal interest rate in the long run. Because inflation is cointegrated with the nominal interest rate, it also is expected to increase by 1 percent in the long run. The main result conveyed by Figure 3 is that inflation and the interest rate already approach their higher long-run levels in the short run. This means that if the increase in X_t^m is interpreted as an increase in the inflation target, the figure suggests that its

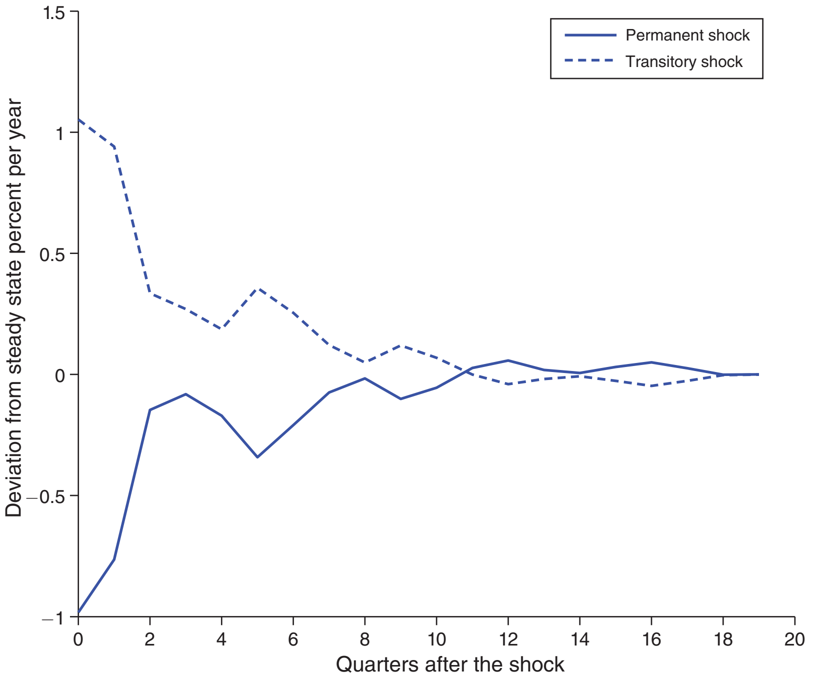

Figure 4. Response of the Real Interest Rate to Permanent and Transitory Monetary Shocks: Empirical Model

Notes: Posterior mean estimates. The real interest rate is defined as $i_t - E_t \pi_{t+1}$.

implementation requires a gradual normalization of the policy rate and results in an immediate reflation. Interestingly, Aruoba and Schorfheide (2011); De Michelis and Iacoviello (2016); and Azevedo, Ritto, and Teles (2019) find a similar result using different empirical methodologies and observables.

On the real side of the economy, the permanent increase in the nominal interest rate does not cause a contraction in aggregate activity. Indeed, output exhibits a transitory expansion.[3] This effect could be the consequence of low real interest rates resulting from the swift reflation of the economy following the permanent interest rate shock. Figure 4 displays with a solid line the response of the real interest rate, defined as $i_t - E_t \pi_{t+1}$, to a permanent interest rate shock. Because of the faster response of inflation relative to that of the nominal interest rate, the real interest rate falls by almost 1 percent on impact and converges to its steady-state level from below, implying that the entire adjustment to a permanent interest rate shock takes place in the context of low real interest rates.

[3] In period 11 the error band narrows to 3 basis points. This is not an uncommon feature of error bands of the type proposed by Sims and Zha (1999) (see, for example, the applications in their paper). It is a reflection of little uncertainty about the position of the impulse response in that period. Additional uncertainty may remain about other features of the impulse response in that period, such as its shape. A similar comment applies for the responses of inflation and output to a temporary monetary shock.

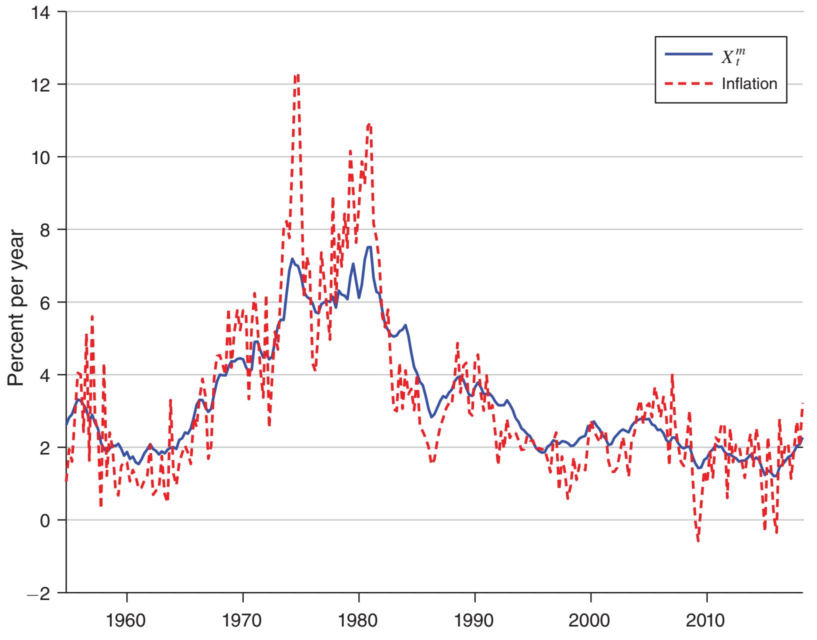

Figure 5. Inflation and Its Permanent Component: Empirical Model

Notes: Quarterly frequency. The inferred path of the permanent component of inflation, X_t^m, was computed by Kalman smoothing and evaluating the empirical model at the posterior mean of the estimated parameter vector. The initial value of X_t^m was normalized to make the average value of X_t^m equal to the average rate of inflation over the sample period, 1954:IV to 2018:II.

The responses of nominal and real variables to a transitory interest rate shock, shown in the right panels of Figure 3, are quite conventional. Both inflation and output fall below trend and remain low for a number of quarters. The real interest rate, whose impulse response is shown with a broken line in Figure 4, increases on impact and remains above its long-run value during the transition, which is in line with the contractionary effect of the transitory increase in the interest rate.

Interestingly, the model does not suffer from the price puzzle, which plagues empirical models with only stationary monetary shocks, pointing to the importance of explicitly distinguishing between temporary and permanent shocks.

A. *Inflation Trends and the Volcker Disinflation*

What does the permanent component of US inflation look like according to the estimated empirical model? Figure 5 displays the actual rate of inflation along with its permanent component, given by the nonstationary monetary shock, X_t^m, over the estimation period (1954:IV to 2018:II). The path of X_t^m resembles the estimate of long-run inflation expectations reported in much of the related empirical literature; see, for example, Cogley and Sargent (2005) and the references cited therein. This result is reassuring because it shows that the short-run effects of temporary and permanent monetary shocks reported in Figure 3 are not based on an estimate of the

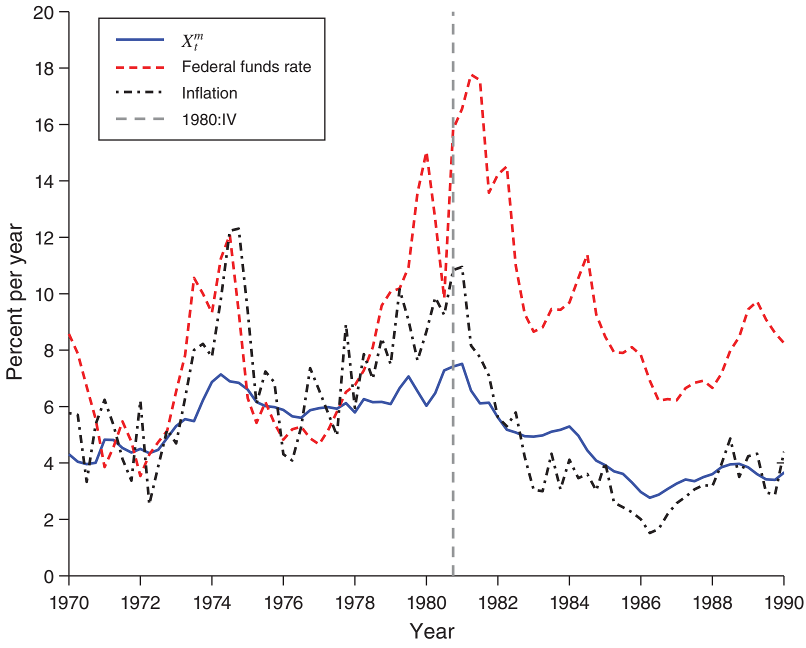

Figure 6. The Volcker Disinflation

Notes: See notes to Figure 5.

permanent component of inflation that is at odds with those obtained in the related literature.

Figure 5 reveals a number of features of the low-frequency drivers of postwar inflation in the United States. First, inflationary factors began to build up much earlier than the oil crisis of 1973. Indeed, the period 1963 to 1972, corresponding to the last seven years in office of Fed Chairman William M. Martin and the first three years of Chairman Arthur F. Burns, were characterized by a continuous increase in the permanent component of inflation, from about 2 percent per year to about 5 percent per year. Second, the high inflation rates associated with the oil crisis of 1973 were not entirely due to nonmonetary shocks. The Fed itself contributed by maintaining X_t^m at the high level it had reached prior to the oil crisis. Third, the figure indicates that the normalization of rates that began in 2015 and put an end to seven years of near-zero nominal rates triggered by the global financial crisis is interpreted by the empirical model as having a significant permanent component.

It is of interest to zoom in on the Volcker era, which arguably represents the largest disinflation episode in the postwar United States. Figure 6 displays the nominal interest rate, the inflation rate, and the permanent monetary shock X_t^m over the period 1970 to 1990. The vertical broken line indicates 1980:IV, which according to Goodfriend and King (2005) represents the beginning of the "deliberate disinflation." The graph suggests that according to the estimated model, the Volcker policy was a combination of a large transitory increase in the policy rate and a gradual decrease in its permanent component. The impulse responses shown in Figure 3 suggest that both of these measures are deflationary. This is consistent with the fact that, as shown in Figure 6, inflation fell faster than its permanent component. Specifically, at the beginning of the stabilization program, 1980:IV, inflation was about 3 percentage points above its permanent component, whereas by 1983, it was already below it, in spite of the fact that the permanent component continued to fall. In fact, one of the most remarkable features of the Volcker disinflation is the speed at which inflation fell. This transition toward low inflation was characterized by depressed economic activity, which is consistent with the enormous magnitude of the hike in the transitory component of the interest rate. According to the empirical model, a decrease in the permanent component of the interest rate would have sufficed to bring about low inflation without unemployment.[4]

B. *Variance Decompositions*

How important are nonstationary monetary shocks? The relevance of the neo-Fisher effect depends on not only whether it can be identified in actual data, which has been the focus of this section thus far, but also whether monetary shocks that change interest rates and inflation in the long run play a significant role in explaining short-run movements in the inflation rate. If nonstationary monetary shocks played a marginal role in explaining cyclical movements in nominal variables, the neo-Fisher effect would just be an interesting curiosity. To shed light on this question, Table 3 displays the variance decomposition of the three variables of interest—output growth, the change in inflation, and the change in the nominal interest rate—predicted by the estimated empirical model. The table shows that the nonstationary monetary shock, X_t^m, explains about 45 percent of the change in inflation, 22 percent of changes in the nominal interest rate, and 9 percent of the growth rate of output. Thus, the empirical model assigns a significant role to nonstationary monetary disturbances, especially in explaining movements in nominal variables. In comparison, the stationary monetary shock, z_t^m, explains a relatively small fraction of movements in the three macroeconomic indicators included in the model.

The permanent monetary shock is also a relevant source of movements in the price level at short and medium time horizons. Figure 7 displays the predicted posterior mean forecast error variance decomposition of output growth, the price level, and the nominal interest rate at horizons 1 to 36 quarters. The nonstationary monetary shock, X_t^m, explains more than 60 percent of movements in the price level at short horizons (1 to 5 quarters) and between 60 and 95 percent at horizons 6 to 36 quarters. By contrast, the transitory monetary shock, z_t^m, explains a small fraction of the forecast error variance of the price level at all horizons.

Taken together, Table 3 and Figure 7 suggest that the shock that generates neo-Fisherian effects, X_t^m, is a relevant driver of nominal variables. More generally, in light of the fact that the majority of studies in monetary economics limit attention to the study of stationary nominal shocks, this result call for devoting more attention

[4] This statement is, of course, subject to the Lucas critique. However, it is confirmed by the optimizing model I study in Section IV.

Table 3—Variance Decomposition: Empirical Model

	Δy_t	$\Delta \pi_t$	Δi_t
Permanent monetary shock, ΔX_t^m	9.1	44.6	21.9
Transitory monetary shock, z_t^m	2.1	6.2	10.9
Permanent nonmonetary shock, ΔX_t	49.8	27.9	13.5
Transitory nonmonetary shock, z_t	39.1	21.4	53.7

Notes: Posterior means. The variables Δy_t, $\Delta \pi_t$, and Δi_t denote output growth, the change in inflation, and the change in the nominal interest rate, respectively.

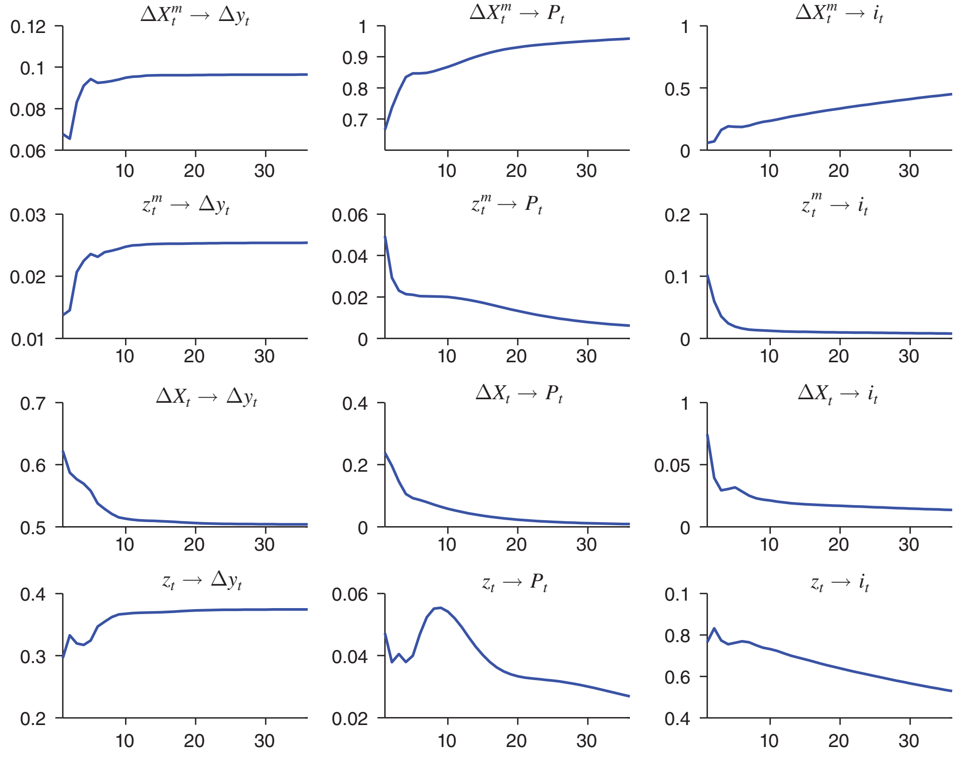

Figure 7. Forecast Error Variance Decomposition Implied by the Empirical Model

Notes: Vertical axes measure shares in percent, and horizontal axes measure forecast horizons in quarters. Forecast error variance shares are posterior mean estimates. Δy_t, P_t, and i_t denote output growth, the price level, and the nominal interest rate, and X_t^m, z_t^m, X_t, and z_t denote the nonstationary monetary shock, the stationary monetary shock, the nonstationary nonmonetary shock, and the stationary nonmonetary shock, respectively.

to understanding the short- and medium-run effects of monetary disturbances that drive the permanent components of inflation and interest rates.

C. *Robustness*

This section considers a number of modifications of the baseline empirical model aimed at gauging the sensitivity of the results. The robustness tests include truncating

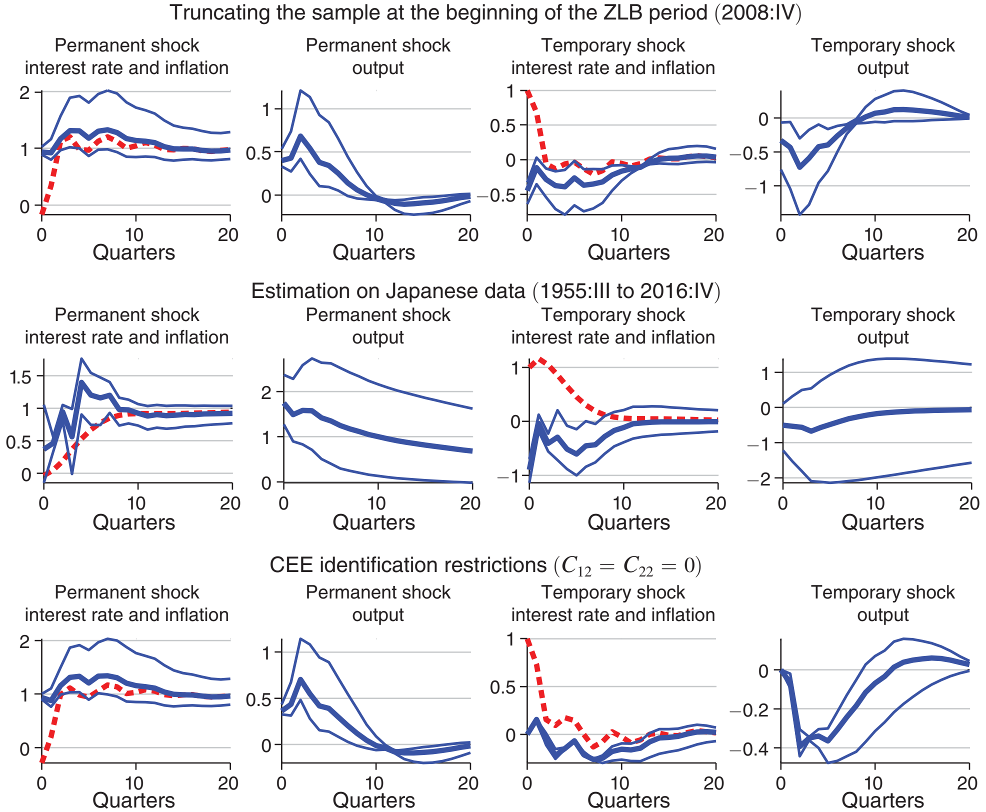

FIGURE 8. ROBUSTNESS CHECKS: EMPIRICAL MODEL

Notes: Thick lines are posterior means. Thick broken lines correspond to the nominal interest rate. Thin lines are 95 percent asymmetric error bands computed using the Sims-Zha (1999) method.

the sample at the beginning of the zero-interest-rate period triggered by the Great Contraction of 2007–2009; estimating the model on Japanese data; identifying the stationary monetary shock à la CEE (2005) by imposing a zero impact effect of a temporary monetary shock on output and inflation, a specification in which the interest-rate-inflation differential is nonstationary; and including the ten-year rate to capture long-run inflationary expectations.

Dropping the Zero Lower Bound Period.—Between 2009 and 2015, the federal funds rate was technically nil, and interest rate policy was said to have hit the zero lower bound (ZLB). The ZLB on nominal rates may introduce nonlinearities that the linear empirical model may not be able to capture. Formulating and estimating a nonlinear model is beyond the scope of this paper. As an imperfect alternative, I estimate the linear model truncating the sample in 2008:IV. The results are shown in the top panels of Figure 8. The impulse responses are qualitatively similar to those obtained with the longer sample.

Estimation on Japanese Data.—As a second robustness check, I estimate the model on Japanese data from 1955:III to 2016:IV. I rely on the results of the previous robustness check in deciding not to truncate the zero-rate period that started in 1995. An additional benefit of keeping the period 1995–2016 is that it might provide valuable information on the effect of permanent monetary shocks, as it involves more than two decades of highly stable rates. The estimated impulse responses appear in the middle row of Figure 8. The figure suggests that the main results obtained using US data carry over to employing Japanese data.

CEE Identification of the Stationary Monetary Shock.—A large number of papers (notably, CEE 2005), identify stationary monetary shocks by assuming that they have a zero impact effect on inflation and output. In the context of the empirical model studied here, this amounts to imposing the restriction

$$C_{12} = C_{22} = 0.$$

The third row of Figure 8 displays the predictions of the empirical model under this identification scheme. The main result of this robustness check is that the predictions of the model are overall in line with those of the baseline specification, which imposes nonpositivity restrictions on the impact effect of a transitory tightening of monetary conditions on output and inflation.

Nonstationary Real Interest Rate.—The baseline model assumes that the policy rate, i_t, and inflation, π_t, are both cointegrated with the permanent monetary shock, X_t^m, with the cointegrating vector $\left[1 - 1\right]$. Under this assumption, i_t and π_t are themselves cointegrated with cointegration vector $\left[1 - 1\right]$. This implies that the real interest rate, $i_t - E_t \pi_{t+1}$, is a stationary variable. Here, I adopt a more flexible specification in which i_t continues to be cointegrated with X_t^m but π_t is cointegrated with αX_t^m, where α is a parameter to be estimated. Under this specification, the interest rate inflation differential, $i_t - \pi_t$, is nonstationary.For this reason, in the vector of observables, I replace it with the change in inflation, $\Delta \pi_t$, which retains its stationarity. The other two observables continue to be Δy_t and Δi_t. I assume that the parameter α has a normal prior distribution with mean 1 and standard deviation 0.15. Its estimated posterior distribution has a mean of 0.9401, a standard deviation of 0.1263, and a 95 percent credible interval of $\left[0.7323, 1.1513\right]$. One cannot reject the hypothesis that the cointegration vector is $\left[1, -1\right]$, as in the baseline case. The top panel of Figure 9 displays the impulse responses of inflation, the policy rate, and output to transitory and permanent monetary shocks. Overall, the key predictions of the baseline model continue to hold under this specification. In particular, the permanent shock generates a quick reflation without output loss, whereas the transitory shock causes a fall in inflation and a contraction in aggregate activity.

Including the Ten-Year Spread.—Intuitively, expanding the baseline model to include a long-maturity rate could help to better discriminate between temporary and more permanent changes in the interest rate, as the latter type of disturbance should be factored in the long rate with a larger loading. Put differently, the addition

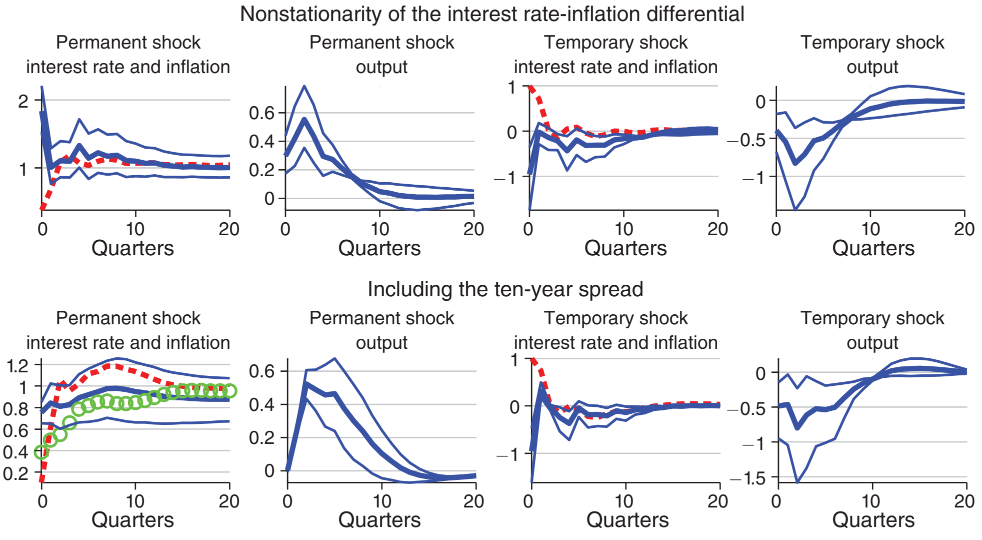

FIGURE 9. ADDITIONAL ROBUSTNESS CHECKS: EMPIRICAL MODEL

Notes: Thick lines are posterior means. Thick broken lines correspond to the nominal interest rate. In the bottom-left panel, the response of the ten-year rate is displayed with circles. Thin lines are 95 percent asymmetric error bands computed using the Sims-Zha (1999) method.

of a long rate should add discipline to the estimation of the permanent monetary shock, as it would be required to be cointegrated with three variables, the inflation rate, the short rate, and the long rate, as opposed to with just the first two variables, as is the case under the baseline formulation.

Figure 10 plots the ten-year rate and the federal funds rate. The ten-year rate is proxied by the Ten-Year Treasury Constant Maturity Rate and is taken from Federal Reserve Economic Data (series GS10). As expected, over the long run, the short and long rates track each other closely. In the short run, the longer rate appears to follow the short rate with some delay.

The empirical model considered here extends the model of Section IIIC to include the ten-year rate, denoted i_t^{10}. Specifically, the unobservable autoregressive system includes the variable $\hat{i}_t^{10} \equiv i_t^{10} - X_t^m$, and the observation equation includes the ten-year spread, $i_t^{10} - i_t$. All other aspects of the model are as in Section IIIC. The bottom panel of Figure 9 displays the impulse responses of output, inflation, the short rate, and the ten-year rate to transitory and permanent monetary shocks. The main predictions of the baseline model extend to the expanded model. In particular, a monetary shock that increases inflation and interest rates in the long run causes an increase in inflation in the short run. As in the raw data, the ten-year rate tracks the short rate with a delay.

Prior Predictions.—Figure 13 in online Appendix C displays prior and posterior responses of inflation, output, and the interest rate to permanent and transitory monetary shocks. The top panel of Figure 14 shows the corresponding responses of the real interest rate. The main results stemming from this exercise are: (i) The posterior

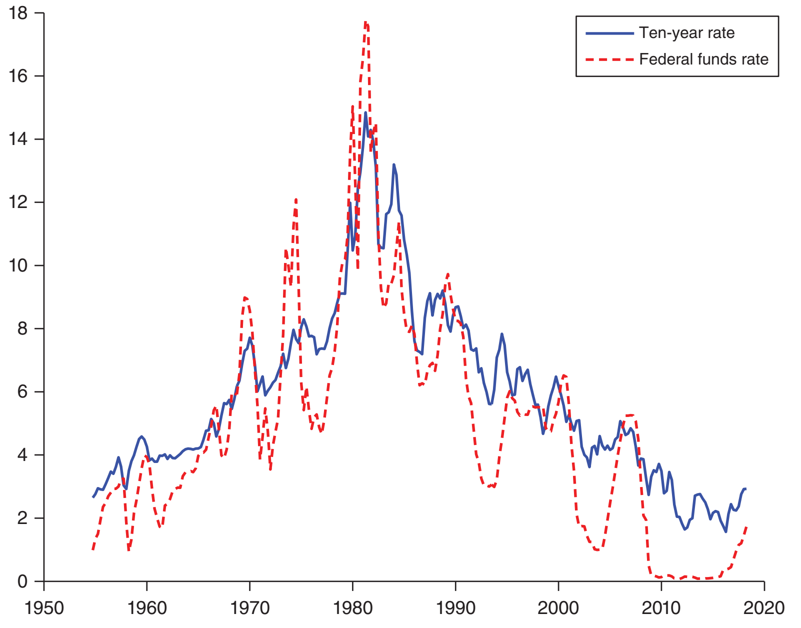

Figure 10. The Ten-Year Rate and the Federal Funds Rate

Note: Quarterly frequency.

Source: See Section IIC.

estimates imply that in response to a permanent monetary shock that increases the interest rate in the long run, the economy reflates much faster than it does under the prior parameterization. (ii) The posterior estimate predicts a transitory expansion in response to a permanent increase in the interest rate, whereas the prior parameterization predicts a mute response. (iii) The posterior estimate predicts a fall in the real interest rate in response to a permanent monetary shock, whereas the prior parameterization predicts a muted response. Interestingly, as shown in Figure 15 and the bottom panel of Figure 14, these results are robust to adopting a CEE-type identification scheme for the transitory monetary shock (see also Section IIIC), in spite of the fact that the prior responses of the nominal interest rate to a temporary monetary shock are quite different under the two schemes.

IV. An Estimated New Keynesian Model with Permanent Trend Inflation Shocks

This section presents an econometric estimation of a small-scale New Keynesian model augmented with a permanent monetary shock (permanent movements in the inflation target) and two temporary monetary shocks, one with high persistence (transitory movements in the inflation target) and one with low persistence. These shocks compete for the data with other monetary and real shocks. The objective of this analysis is not theoretical in nature. A number of papers cited in the introduction have shown that in models of this type, permanent and stationary but persistent changes in the inflation target are implemented via rising interest rates and inflation in the short run. The goal of this section is twofold. One is to ascertain from the perspective of a standard New Keynesian DSGE model the importance of the monetary shocks that produce neo-Fisherian effects. The other is to establish whether these effects stem primarily from stationary or nonstationary movements in the inflation target as formulated in, for example, Garín, Lester, and Sims (2018). The second objective cannot be implemented with the semistructural model studied thus far. The optimizing nature of the DSGE model, by contrast, makes this estimation possible.

The model features price stickiness and habit formation and is driven by four real shocks in addition to the aforementioned three monetary shocks: permanent and transitory productivity shocks, a preference shock, and a labor supply shock. This section presents the main building blocks of the model. Online Appendix B offers a detailed derivation of the equilibrium conditions.

The economy is populated by households with preferences defined over streams of consumption and labor effort and exhibiting external habit formation. The household's lifetime utility function is

$$E_0 \sum_{t=0}^{\infty} \beta^t e^{\xi_t} \left\{ \frac{\left[\left(C_t - \delta \tilde{C}_{t-1} \right) \left(1 - e^{\theta_t} h_t \right)^{\chi} \right]^{1-\sigma} - 1}{1-\sigma} \right\}, \tag{5}$$

where C_t denotes consumption, $\tilde{C}_t$ denotes the cross-sectional average of consumption, h_t denotes hours worked, ξ_t is a preference shock, θ_t is a labor supply shock, and $\beta, \delta \in (0,1)$ and $\sigma, \chi > 0$ are parameters.

Households are subject to the budget constraint

$$P_t C_t + \frac{B_{t+1}}{1+I_t} + T_t = B_t + W_t h_t + \Phi_t, \tag{6}$$

where P_t denotes the nominal price of consumption, B_{t+1} denotes a nominal bond purchased in t and paying the nominal interest rate I_t in $t+1$, T_t denotes nominal lump-sum taxes, W_t denotes the nominal wage rate, and Φ_t denotes nominal profits received from firms.

The consumption good C_t is assumed to be a composite of a continuum of varieties C_{it} indexed by $i \in [0,1]$ with aggregation technology $C_t = \left[\int_0^1 C_{it}^{1-1/\eta} di \right]^{\frac{1}{1-1/\eta}}$, where the parameter $\eta > 0$ denotes the elasticity of substitution across varieties.

The firm producing variety i operates in a monopolistically competitive market and faces quadratic price adjustment costs à la Rotemberg (1982). The production technology uses labor and is buffeted by stationary and nonstationary productivity shocks. Specifically, output of variety i is given by

$$Y_{it} = e^{z_t} X_t h_{it}^{\alpha}, \tag{7}$$

where Y_{it} denotes output of variety i in period t, h_{it} denotes labor input used in the production of variety i, and z_t and X_t are stationary and nonstationary productivity shocks, respectively. The growth rate of the nonstationary productivity shock, $g_t \equiv \ln(X_t/X_{t-1})$, is assumed to be a stationary random variable. The expected present discounted value of real profits of the firm producing variety i expressed in units of consumption is given by

$$E_0 \sum_{t=0}^{\infty} q_t \left[\frac{P_{it}}{P_t} C_{it} - \frac{W_t}{P_t} h_{it} - \frac{\phi}{2} X_t \left(\frac{P_{it}/P_{it-1}}{1+\tilde{\Pi}_t} - 1 \right)^2 \right], \tag{8}$$

where $1+\tilde{\Pi}_t = (1+\tilde{\Pi}_{t-1})^{\gamma_m}(1+\Pi_t)^{1-\gamma_m}$ denotes the average level of inflation around which price adjustment costs are defined and $\Pi_t \equiv P_t/P_{t-1} - 1$ denotes the inflation rate. The parameter $\phi > 0$ governs the degree of price stickiness, and the parameter $\gamma_m \in [0,1]$ the backward-looking component of the inflation measure at which price adjustment costs are centered. Both parameters are estimated. Allowing for a backward-looking component in firms' price-setting behavior is in order in the present context because, as pointed out by Garín, Lester, and Sims (2018), the larger this component is, the less likely it will be that stationary but persistent movements in the inflation target are implemented with rising interest rates and inflation in the short run. The variable $q_t \equiv \beta^t \frac{\Lambda_t}{\Lambda_0}$ denotes a pricing kernel reflecting the assumption that profits belong to households. The price adjustment cost in the profit equation (8) is scaled by the output trend X_t to keep nominal rigidity from vanishing along the balanced growth path.

The monetary authority follows a Taylor-type interest rate feedback rule with smoothing, as follows:

$$\frac{1+I_t}{\Gamma_t} = \left[A \left(\frac{1+\Pi_t}{\Gamma_t} \right)^{\alpha_\pi} \left(\frac{Y_t}{X_t} \right)^{\alpha_y} \right]^{1-\gamma_I} \left(\frac{1+I_{t-1}}{\Gamma_{t-1}} \right)^{\gamma_I} e^{z_t^m},$$

where Y_t denotes aggregate output; z_t^m denotes a stationary interest rate shock; Γ_t is the inflation target; and A, α_π, α_y, and $\gamma_I \in [0,1)$ are parameters. The inflation target is assumed to have a permanent component denoted X_t^m and a transitory component denoted z_t^{m2}. Formally,

$$\Gamma_t = X_t^m e^{z_t^{m2}}.$$

The growth rate of the permanent component of the inflation target, $g_t^m \equiv \ln(X_t^m/X_{t-1}^m)$, is assumed to be stationary. Up to first order, the stationary component of the inflation target can be observationally equivalent to a standard monetary shock with nonzero serial correlation. It is therefore in order to comment on the identification of z_t^m and z_t^{m2}. The distinction of these two stationary monetary shocks is achieved by imposing restrictions on the prior distribution of their serial correlations. Specifically, the serial correlation of z_t^m is assumed to have a prior mean of 0.3, and the serial correlation of z_t^{m2} a prior mean of 0.7. (Interest rate smoothing—that is, an estimate of γ_I significantly different from zero—adds an additional identification channel for z_t^{m2}.)

Table 4—Calibrated Parameters in the New Keynesian Model

Parameter	Value	Description
β	0.9982	Subjective discount factor
σ	2	Inverse of intertemp. elast. subst.
η	6	Intratemporal elast. of subst.
α	0.75	Labor semielast. of output
g	0.004131	Mean output growth rate
θ	0.4055	Preference parameter
χ	0.625	Preference parameter

Note: The time unit is one quarter.

Government consumption is assumed to be nil at all times, and fiscal policy is assumed to be Ricardian.

The seven structural shocks driving the economy, ξ_t, θ_t, z_t, g_t, z_t^m, z_t^{m2}, and g_t^m, are assumed to follow AR(1) processes of the form $x_t = \rho_x x_{t-1} + \epsilon_t^x$, for $x = \xi, \theta, z, g, z^m, z^{2m}, g^m$.

As in much of the DSGE literature, I estimate a subset of the parameters of the model and calibrate the remaining ones using standard values in business cycle analysis. The set of estimated parameters includes those that play a central role in determining the model's implied short-run dynamics, such as those defining price adjustment costs, habit formation, monetary policy, and the stochastic properties of the underlying sources of uncertainty. Table 4 displays the values assigned to the calibrated parameters. I set the subjective discount factor, β, equal to 0.9982, which implies a growth-adjusted discount factor, $\beta e^{-\sigma g}$, equal to 0.99; the reciprocal of the intertemporal elasticity of substitution, σ, to 2; the intratemporal elasticity of substitution across varieties of intermediate goods, η, to 6; the labor elasticity of the production function, α, to 0.75; the unconditional mean of per capita output growth, g, to 0.004131 (1.65 percent per year), which matches the average growth rate of real GDP per capita in the United States over the estimation period (1954:IV to 2018:II); and the parameters θ and χ to ensure, given all other parameter values, that in the steady state, households allocate one-third of their time to work, $h = 1/3$, and a unit Frisch elasticity of labor supply, $\left(1 - e^{\theta} h\right) / \left(e^{\theta} h\right) = 1$ (Galí 2008).

The remaining parameters of the model are estimated using Bayesian techniques and the same observables as in the estimation of the semistructural model of Section II—namely, per capita output growth, the interest rate–inflation differential, and the change in the nominal interest rate. Table 5 displays summary statistics of the prior and posterior distributions of the estimated parameters. Draws from the posterior distribution are based on a random walk Metropolis-Hastings MCMC chain of length 1 million after discarding 100,000 burn-in draws. Most parameters are estimated with significant uncertainty, a feature that is common in estimates of small-scale New Keynesian models (Ireland 2007). Nonetheless, the data speak with a strong voice on the parameters ϕ and δ, respectively governing price stickiness and habit formation, which are key determinants of the propagation of nominal and real shocks.

Figure 11 displays the estimated impulse responses of inflation, the policy rate, and output to inflation target shocks (X_t^m and z_t^{m2}) and interest rate shocks

Table 5—Prior and Posterior Parameter Distributions: New Keynesian Model

Parameter	Prior distribution			Posterior distribution			
	Distribution	Mean	SD	Mean	SD	5 percent	95 percent
ϕ	Gamma	50	20	146	31.9	96.8	201
α_π	Gamma	1.5	0.25	2.32	0.221	1.98	2.7
α_y	Gamma	0.125	0.1	0.188	0.123	0.0336	0.422
γ_m	Uniform	0.5	0.289	0.606	0.0762	0.475	0.724
γ_I	Uniform	0.5	0.289	0.242	0.142	0.053	0.517
δ	Uniform	0.5	0.289	0.258	0.0531	0.173	0.348
ρ_ξ	Beta	0.7	0.2	0.915	0.0234	0.874	0.95
ρ_θ	Beta	0.7	0.2	0.708	0.21	0.317	0.98
ρ_z	Beta	0.7	0.2	0.7	0.214	0.302	0.978
ρ_g	Beta	0.3	0.2	0.221	0.108	0.0557	0.41
ρ_{gm}	Beta	0.3	0.2	0.248	0.166	0.0295	0.562
ρ_{zm}	Beta	0.3	0.2	0.306	0.184	0.0526	0.654
ρ_{zm2}	Beta	0.7	0.2	0.796	0.205	0.33	0.975
σ_ξ	Gamma	0.01	0.01	0.0287	0.00602	0.0212	0.0398
σ_θ	Gamma	0.01	0.01	0.00164	0.00138	0.000115	0.00435
σ_z	Gamma	0.01	0.01	0.00122	0.000974	8.66e-05	0.00312
σ_g	Gamma	0.01	0.01	0.00758	0.000944	0.00593	0.00905
σ_{gm}	Gamma	0.0025	0.0025	0.000848	0.000474	8.48e-05	0.00159
σ_{zm}	Gamma	0.0025	0.0025	0.000832	0.000465	7.96e-05	0.00152
σ_{zm2}	Gamma	0.0025	0.0025	0.00131	0.000733	0.000138	0.00248
R_{11}	Gamma	3.78e-06	2.18e-06	4.46e-06	2.59e-06	1.22e-06	9.46e-06
R_{22}	Gamma	2.08e-06	1.2e-06	4.55e-06	4.88e-07	3.79e-06	5.4e-06
R_{33}	Gamma	2.36e-07	1.36e-07	1.74e-07	9.95e-08	4.82e-08	3.62e-07

Notes: The time unit is one quarter. Growth rates and log deviations from trend are expressed in per one (1 percent is denoted 0.01).

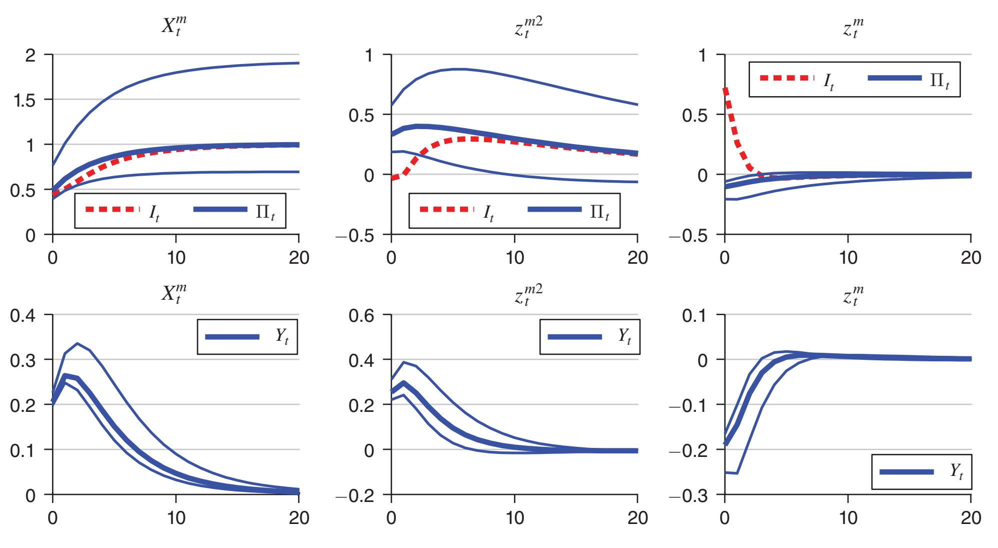

Figure 11. Estimated Impulse Responses to Inflation Target and Interest Rate Shocks in the New Keynesian Model

Notes: Impulse responses are posterior mean estimates. Inflation, Π_t, and the nominal interest rate, I_t, are deviations from preshock levels and expressed in percentage points per year. Output, Y_t, is measured in percent deviations from trend. Thin lines represent 95 percent credible intervals for inflation (top panels) and output (bottom panels). Asymmetric error bands are computed using the Sims-Zha (1999) method.

(z_t^m) implied by the estimated New Keynesian model. The main message conveyed by the figure is that qualitatively the responses implied by the New Keynesian model concur with those implied by the empirical model of Sections II and III. In the estimated New Keynesian model, a permanent increase in the inflation target, X_t^m, is implemented with a gradual increase in the nominal interest rate, which reaches its higher long-run level in about ten quarters. In response to this policy innovation, inflation increases monotonically to its new steady-state value, without loss of aggregate activity. Similarly, an increase in the transitory component of the inflation target, z_t^{m2}, causes rising interest rates, an elevation in the rate of inflation, and no contraction in output.

The estimated response of inflation and the interest rate to a stationary increase in the inflation target provides econometric support to the theoretical finding of Garín, Lester, and Sims (2018) that stationary trend shocks can produce neo-Fisherian effects if sufficiently persistent. Although ρ_{zm2} is estimated with significant uncertainty, the data pick a mean posterior value higher than its prior counterpart (0.8 versus 0.7). By contrast, the standard transitory interest rate shock, z_t^m, is estimated to cause a fall in inflation and a contraction in aggregate activity.

Figure 11 shows that in response to either a permanent or a transitory but persistent increase in the inflation target, inflation not only begins to increase immediately but does so at a rate faster than the nominal interest rate. As a result, the real interest rate falls, as shown in Figure 12. By contrast, a short-lived increase in the nominal interest rate causes a fall in inflation and an increase in the real interest rate. A natural question is why inflation moves faster than the interest rate in the short run when the monetary shock is expected to be permanent or transitory but persistent. The answer has to do with the presence of nominal rigidities and with the way the central bank conducts monetary policy. In response to an increase in the inflation target, the central bank raises the short-run policy rate quickly but gradually. At the same time, firms know that by the classic Fisher effect, the consumer price level and the nominal wage will increase down the road. They therefore realize that if they don't follow suit, they will face ever-increasing losses as time goes by, since they would sell their product increasingly cheaply relative to other firms while facing elevated labor costs. Since firms face quadratic costs of adjusting prices, they find it optimal to begin increasing their price immediately. And since all firms do the same, inflation itself begins to increase as soon as the shock occurs.

The central contribution of this section is to ascertain the importance of the shocks that have neo-Fisherian effects, X_t^m and z_t^{m2}, in explaining movements in the inflation rate. Table 6 displays this information. The permanent monetary shock, X_t^m, explains more than 30 percent of the variance of changes in the rate of inflation. Thus, like the empirical model, the New Keynesian model assigns a significant role to permanent innovations in monetary policy. Transitory movements in the inflation target, embodied in the shock z_t^{m2}, explain 22 percent of changes in the rate of inflation. Thus, trend inflation shocks $(X_t^m$ and $z_t^{m2})$ jointly explain more than 50 percent of the variance of changes in the inflation rate. Also, as in the empirical model, in the New Keynesian model, the stationary interest rate shock, z_t^m, accounts for a

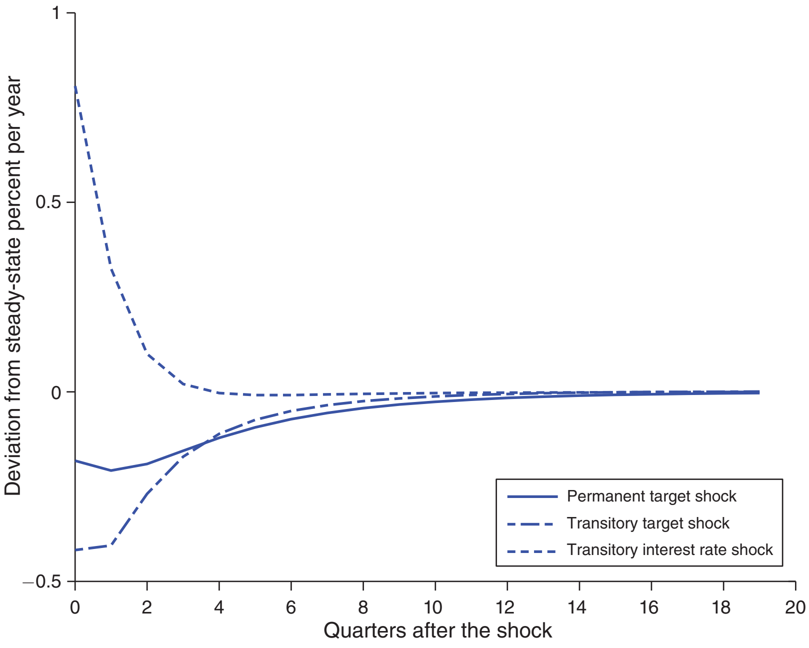

Figure 12. Estimated Response of the Real Interest Rate to Inflation Target and Interest Rate Shocks in the New Keynesian Model

Notes: Posterior mean estimates. The real interest rate is defined as $i_t - E_t \pi_{t+1}$.

Table 6— Variance Decomposition: New Keynesian Model

	Δy_t	$\Delta \pi_t$	Δi_t
Permanent trend inflation shock, g_t^m	2.4	30.1	7.6
Transitory trend inflation shock, z_t^{m2}	4.3	22.2	5.1
Transitory interest rate shock, z_t^m	1.2	1.2	14.2
Permanent productivity shock, g_t	79.5	0.8	1.6
Transitory productivity shock, z_t	0.5	2.8	2.6
Preference shock, ξ_t	11.5	40.0	66.3
Labor supply shock, θ_t	0.6	2.8	2.6

Notes: Posterior means. The variables Δy_t, $\Delta \pi_t$, and Δi_t denote output growth, the change in inflation, and the change in the nominal interest rate, respectively.

relatively small share of movements in the rate of inflation. Taken together, these results indicate that monetary shocks that induce neo-Fisherian dynamics appear to have a significance presence in the data.

V. Conclusion

Discussions of how monetary policy can lift an economy out of chronic below-target inflation are almost always based on the logic of how transitory interest rate shocks affect real and nominal variables. Nowadays, there is little theoretical or empirical controversy around how transitory monetary shocks transmit to the rest of the economy: an increase in the nominal interest rate causes an increase in the real interest rate, which puts downward pressure on both aggregate activity and price growth. Within this logic, a central bank trying to reflate a low-inflation economy will tend to set interest rates as low as possible. This policy is effective as long as the cut in interest rates is expected to be transitory.

The question is what happens when the low interest rate policy has been in place for a decade or more and agents come to expect that low rates will continue to be maintained over the indefinite future—as in Japan post-1995, the eurozone post-2008, or, as it seems, the United States post-COVID-19 pandemic. The available evidence shows that at some point, these economies find themselves with zero or negative nominal rates and with the low-inflation problem not going away. One interpretation of what happens at this point is that the situation perpetuates: the monetary authority keeps the interest rate low because inflation is still below target (the temporary-interest-rate-shock logic), and inflation is low because the interest rate has been low for a long period of time (the classic Fisher effect).

In this paper, I provide empirical evidence drawn from an empirical and an optimizing model in favor of the hypothesis that a gradual and permanent increase in the nominal interest rate leads to a quick and monotonic adjustment of inflation to a permanently higher level, low real interest rates, and no output loss. Put differently, implementing an increase in the inflation target requires gradually rising rates and causes a rising path of inflation.

Taken together, the findings reported in this paper are consistent with the neo-Fisherian prediction that a credible announcement of a gradual return of the nominal interest rate from the vicinity of zero to historically normal levels can achieve a swift reflation of the economy with sustained levels of economic activity.

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