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

By Martín Uribe

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
their higher long-run levels within a year. Furthermore, the adjustment to a perma-
nent 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 perma-
nent and stationary inflation target shocks compete with standard transitory mon-
etary 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 res-
olution 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 main-
tains 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 mon-
etary 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 mone-
tary 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 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
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.

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

<table>
<thead>
<tr>
<th>Transitory shock</th>
<th>Long run effect</th>
<th>Short run effect</th>
</tr>
</thead>
<tbody>
<tr>
<td>Permanent shock</td>
<td>↑</td>
<td>↑</td>
</tr>
</tbody>
</table>

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 $\pi_t$ 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 $y_t$, $\pi_t$, 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 non-

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

Note: Quarterly frequency.
Source: See Section IIIC
stationary 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 \( X^m_t \). I then define the following vector of stationary variables:

\[
\begin{bmatrix}
\hat{y}_t \\
\hat{\pi}_t \\
\hat{i}_t
\end{bmatrix}
\equiv
\begin{bmatrix}
y_t - X_t \\
\pi_t - X^m_t \\
i_t - X^m_t
\end{bmatrix}.
\]

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 \( [\hat{y}_t \ \hat{\pi}_t \ \hat{i}_t]^\prime \) is assumed to take the autoregressive form

\[
\begin{bmatrix}
\hat{y}_t \\
\hat{\pi}_t \\
\hat{i}_t
\end{bmatrix} = \sum_{i=1}^{4} B_i \begin{bmatrix}
\hat{y}_{t-i} \\
\hat{\pi}_{t-i} \\
\hat{i}_{t-i}
\end{bmatrix} + C \begin{bmatrix}
\Delta X^m_t \\
\Delta X^m_t \\
z_t
\end{bmatrix},
\]

where \( \Delta X^m_t \equiv X^m_t - X^m_{t-1} \), \( \Delta X_t \equiv X_t - X_{t-1} \), 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

\[
\begin{bmatrix}
\Delta X^m_{t+1} \\
z^m_{t+1} \\
\Delta X^m_{t+1} \\
z_{t+1}
\end{bmatrix} = \rho \begin{bmatrix}
\Delta X^m_t \\
z^m_t \\
\Delta X_t \\
z_t
\end{bmatrix} + \psi \begin{bmatrix}
\epsilon^1_{t+1} \\
\epsilon^2_{t+1} \\
\epsilon^3_{t+1} \\
\epsilon^4_{t+1}
\end{bmatrix},
\]

where \( \rho \) and \( \psi \) are diagonal matrices of coefficients to be estimated and \( \epsilon^i_t \) 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 \( X^m_t \). In addition, to identify the transitory monetary shock, \( z^m_t \), 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 $z_{tm}$ from $z_t$ are imposed via restrictions on the prior distributions of the elements of $C$ and $\rho$. 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, $\Delta y_t$; the change in the nominal interest rate, $\Delta i_t$; 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):

\[
\begin{align*}
\Delta y_t &= \hat{y}_t - \hat{y}_{t-1} + \Delta X_t, \\
r_t &= \hat{i}_t - \hat{\pi}_t, \\
\Delta i_t &= \hat{i}_t - \hat{i}_{t-1} + \Delta X_t^m.
\end{align*}
\]

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

\[
o_t = \begin{bmatrix} \Delta y_t \\ r_t \\ \Delta i_t \end{bmatrix} + \mu_t,
\]

where $\mu_t$ is a vector of measurement errors distributed i.i.d. $\mathcal{N}(\emptyset, R)$ 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...
errors are restricted to explain no more than 10 percent of the variance of the observ-
ables. 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 \( \pi_t \) 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 \( B_i, i = 1, \ldots, 4, C, \rho, \) and \( \psi \)
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
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

<table>
<thead>
<tr>
<th>Parameter Description</th>
<th>Distribution</th>
<th>Mean</th>
<th>Standard Deviation</th>
</tr>
</thead>
<tbody>
<tr>
<td>Diagonal elements of $B_1$</td>
<td>Normal</td>
<td>0.95</td>
<td>0.5</td>
</tr>
<tr>
<td>All other elements of $B_i$, $i = 1, 2, 3, 4$</td>
<td>Normal</td>
<td>0</td>
<td>0.25</td>
</tr>
<tr>
<td>$C_{21}, C_{31}$</td>
<td>Normal</td>
<td>$-1$</td>
<td>1</td>
</tr>
<tr>
<td>$-C_{12}, -C_{22}$</td>
<td>Normal</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>All other estimated elements of $C$</td>
<td>Gamma</td>
<td>1</td>
<td>1</td>
</tr>
<tr>
<td>$\psi_{ii}, i = 1, 2, 3, 4$</td>
<td>Beta</td>
<td>0.3</td>
<td>0.2</td>
</tr>
<tr>
<td>$\rho_{ii}, i = 1, 2, 4$</td>
<td>Beta</td>
<td>0.7</td>
<td>0.2</td>
</tr>
<tr>
<td>$R_{ii}, i = 1, 2, 3$</td>
<td>Uniform $\left[0, \frac{\text{var}(\alpha)}{10}\right]$</td>
<td>$\frac{\text{var}(\alpha)}{10 \times 2}$</td>
<td>$\frac{\text{var}(\alpha)}{10 \times \sqrt{12}}$</td>
</tr>
</tbody>
</table>

²One 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 $y_t = X_t + z + \alpha k_t + (1 - \alpha) h_t$ 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...
\( \psi_{ii} \), 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, \( \rho_{ii} \), \( i = 1, \ldots, 4 \). 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 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 \( \rho \) 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 \( \rho \) 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...
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
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. 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.
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.

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
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

Figure 6. The Volcker Disinflation

Notes: See notes to Figure 5.
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.
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

---

**Table 3—Variance Decomposition: Empirical Model**

<table>
<thead>
<tr>
<th></th>
<th>$\Delta y_t$</th>
<th>$\Delta \pi_t$</th>
<th>$\Delta i_t$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Permanent monetary shock, $\Delta X_t^m$</td>
<td>9.1</td>
<td>44.6</td>
<td>21.9</td>
</tr>
<tr>
<td>Transitory monetary shock, $z_t^m$</td>
<td>2.1</td>
<td>6.2</td>
<td>10.9</td>
</tr>
<tr>
<td>Permanent nonmonetary shock, $\Delta X_t$</td>
<td>49.8</td>
<td>27.9</td>
<td>13.5</td>
</tr>
<tr>
<td>Transitory nonmonetary shock, $z_t$</td>
<td>39.1</td>
<td>21.4</td>
<td>53.7</td>
</tr>
</tbody>
</table>

*Notes: Posterior means. The variables $\Delta y_t$, $\Delta \pi_t$, and $\Delta i_t$ 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. $\Delta y_t$, $P_t$, 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.*
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, \( \pi_t \), are both cointegrated with the permanent monetary shock, \( X_{tm} \), with the cointegrating vector \([1 \quad -1]\). Under this assumption, \( i_t \) and \( \pi_t \) are themselves cointegrated with cointegration vector \([1 \quad -1]\). This implies that the real interest rate, \( i_t - E_{t+1} \pi_t \), is a stationary variable. Here, I adopt a more flexible specification in which \( i_t \) continues to be cointegrated with \( X_{tm} \) but \( \pi_t \) is cointegrated with \( \alpha X_{tm} \), where \( \alpha \) 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 \( \Delta y_t \) and \( \Delta i_t \). I assume that the parameter \( \alpha \) 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 \([0.7323, 1.1513]\). One cannot reject the hypothesis that the cointegration vector is \([1 \quad -1]\), 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
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 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
estimates imply that in response to a permanent monetary shock that increases the interest rate in the long run, the economy reducts 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

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

Note: Quarterly frequency.
Source: See Section IIC.
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

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

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

Households are subject to the budget constraint

\[ P_t C_t + \frac{B_{t+1}}{1 + I_t} + T_t = B_t + W_t h_t + \Phi_t, \]

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 \( \Phi_t \) 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 \( C_t = \left[ \int_0^1 C_{it}^{-1/\eta} \, dt \right]^{1/1/\eta} \), 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

\[ Y_{it} = e^{\tilde{z}_t} X_t h_{it}^\alpha, \]

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

\[
E_0 \sum_{t=0}^{\infty} q_t \left[ \frac{P_{it}}{P_t} C_{it} - W_t h_{it} - \frac{\phi}{2} X_t \left( \frac{P_{it}}{P_{it-1}} - 1 \right) \right]^2,
\]

where \( 1 + \tilde{\Pi}_t = (1 + \tilde{\Pi}_{t-1})^{1-\gamma_m} \) denotes the average level of inflation around which price adjustment costs are defined and \( \Gamma_t \equiv P_t/P_{t-1} - 1 \) 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 \( q_t \equiv \beta_t \Lambda_t \) 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:

\[
\frac{1 + I_t}{\Gamma_t} = A \left( \frac{1 + \Pi_t}{\Gamma_t} \right)^{\alpha_z} \left( \frac{Y_t}{X_t} \right)^{\alpha_y} \left( \frac{1 + I_{t-1}}{\Gamma_{t-1}} \right)^{\gamma_I} e^{z_{tm}},
\]

where \( Y_t \) denotes aggregate output; \( z_{tm} \) denotes a stationary interest rate shock; \( \Gamma_t \) is the inflation target; and \( A, \alpha_z, \alpha_y, \) and \( \gamma_I \in [0, 1] \) are parameters. The inflation target is assumed to have a permanent component denoted \( X_{tm} \) and a transitory component denoted \( z_{tm}^2 \). Formally,

\[
\Gamma_t = X_{tm} e^{z_{tm}^2}.
\]

The growth rate of the permanent component of the inflation target, \( g_{tm}^m \equiv \ln(X_{tm}/X_{tm-1}) \), 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 \( z_{tm} \) and \( z_{tm}^2 \). 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 \( z_{tm} \) is assumed to have a prior mean of 0.3, and the serial correlation of \( z_{tm}^2 \) a prior mean of 0.7. (Interest rate smoothing—that is, an estimate of \( \gamma_I \) significantly different from zero—adds an additional identification channel for \( z_{tm}^2 \).)
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, $\xi_t$, $\theta_t$, $z_t$, $g_t$, $z_t^m$, $z_t^{2m}$, and $g_t^m$, are assumed to follow AR(1) processes of the form $x_t = \rho x_{t-1} + \epsilon_t$, for $x = \xi_t, \theta_t, z_t, g_t, z_t^m, z_t^{2m}, g_t^m$.

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, $\beta$, 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, $\sigma$, to 2; the intratemporal elasticity of substitution across varieties of intermediate goods, $\eta$, to 6; the labor elasticity of the production function, $\alpha$, 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 $\theta$ and $\chi$ 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, $(1 - e^{\theta h}) / (e^{\theta h}) = 1$ (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 $\phi$ and $\delta$, 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

<table>
<thead>
<tr>
<th>Parameter</th>
<th>Prior distribution</th>
<th>Posterior distribution</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Distribution</td>
<td>Mean</td>
</tr>
<tr>
<td>$\phi$</td>
<td>Gamma 50 20</td>
<td>146</td>
</tr>
<tr>
<td>$\alpha_\pi$</td>
<td>Gamma 1.5 0.25</td>
<td>2.32</td>
</tr>
<tr>
<td>$\alpha_y$</td>
<td>Gamma 0.125 0.1</td>
<td>0.188</td>
</tr>
<tr>
<td>$\gamma_m$</td>
<td>Uniform 0.5 0.289</td>
<td>0.606</td>
</tr>
<tr>
<td>$\gamma_y$</td>
<td>Uniform 0.5 0.289</td>
<td>0.242</td>
</tr>
<tr>
<td>$\delta$</td>
<td>Uniform 0.5 0.289</td>
<td>0.258</td>
</tr>
<tr>
<td>$\rho_\xi$</td>
<td>Beta 0.7 0.2</td>
<td>0.915</td>
</tr>
<tr>
<td>$\rho_\theta$</td>
<td>Beta 0.7 0.2</td>
<td>0.708</td>
</tr>
<tr>
<td>$\rho_y$</td>
<td>Beta 0.7 0.2</td>
<td>0.7</td>
</tr>
<tr>
<td>$\rho_m$</td>
<td>Beta 0.3 0.2</td>
<td>0.221</td>
</tr>
<tr>
<td>$\rho_{gm}$</td>
<td>Beta 0.3 0.2</td>
<td>0.248</td>
</tr>
<tr>
<td>$\rho_{zm}$</td>
<td>Beta 0.3 0.2</td>
<td>0.306</td>
</tr>
<tr>
<td>$\rho_{zm2}$</td>
<td>Beta 0.7 0.2</td>
<td>0.796</td>
</tr>
<tr>
<td>$\sigma_\xi$</td>
<td>Gamma 0.01 0.01</td>
<td>0.0287</td>
</tr>
<tr>
<td>$\sigma_\theta$</td>
<td>Gamma 0.01 0.01</td>
<td>0.00164</td>
</tr>
<tr>
<td>$\sigma_y$</td>
<td>Gamma 0.01 0.01</td>
<td>0.00122</td>
</tr>
<tr>
<td>$\sigma_\theta$</td>
<td>Gamma 0.01 0.01</td>
<td>0.00758</td>
</tr>
<tr>
<td>$\sigma_{gm}$</td>
<td>Gamma 0.0025 0.0025</td>
<td>0.000848</td>
</tr>
<tr>
<td>$\sigma_{zm}$</td>
<td>Gamma 0.0025 0.0025</td>
<td>0.000832</td>
</tr>
<tr>
<td>$\sigma_{zm2}$</td>
<td>Gamma 0.0025 0.0025</td>
<td>0.00131</td>
</tr>
<tr>
<td>$\sigma_{zm2}$</td>
<td>Gamma 3.78e-06 2.18e-06</td>
<td>4.46e-06</td>
</tr>
<tr>
<td>$\sigma_{zm}$</td>
<td>Gamma 2.08e-06 1.2e-06</td>
<td>4.55e-06</td>
</tr>
<tr>
<td>$\sigma_{zm}$</td>
<td>Gamma 2.36e-07 1.36e-07</td>
<td>1.74e-07</td>
</tr>
</tbody>
</table>

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, $\Pi_t$, 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.
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 $\rho_{zm2}$ 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
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 opti-
mizing 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.

REFERENCES

Aruoba, S. Borag˘an, and Frank Schorfheide. 2011. “Sticky Prices versus Monetary Frictions: An Esti-
Long is the Long Run?” Banco de Portugal Working Paper 11.
August 6, 2018).

Christiano, Lawrence J., Martin Eichenbaum, and Charles L. Evans. 2005. “Nominal Rigidities and


Quiet Inflation at the Zero Bound.” In NBER Macroeconomics Annual, Vol. 32, edited by Martin


DeMichelis, Andrea, and Matteo Iacoviello. 2016. “Raising an Inflation Target: The Japanese Experi-

Erceg, Christopher J., and Andrew T. Levin. 2003. “Imperfect Credibility and Inflation Persistence?”

Fève, Patrick, Julien Matheron, and Jean-Guillaume Sahuc. 2010. “Disinflation Shocks in the


