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# THE MACROECONOMIC CONSEQUENCES OF NATURAL RATE SHOCKS: AN EMPIRICAL INVESTIGATION

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**ABSTRACT**

Much of the empirical literature on the natural rate of interest has focused on estimating its path. This paper addresses the question of how exogenous movements in the natural rate of interest affect aggregate activity and inflation in the short and long runs. To this end it proposes a semi-structural model of output, inflation, and the policy interest rate inspired by the DSGE literature but with fewer identification and cross-equation restrictions. It then estimates it on U.S. data over the period 1900 to 2021. We find that a permanent decline in the natural rate of interest has a large negative effect on the trend of output and is contractionary and deflationary in the short run. When the economy is constrained by the zero lower bound (ZLB), these results are consistent with the secular stagnation hypothesis. However, we find that negative natural rate shocks depress the trend of output even when the economy is away from the ZLB. Thus, the results of this paper call for a more general theory of the trend effects of natural rate shocks.

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# 1 Introduction

The real rate of interest has displayed a persistent decline over the past decades. A substantial amount of empirical research has been devoted to ascertaining whether this phenomenon corresponds to a fall in the natural rate of interest, understood as the permanent component of the short-term real interest rate. Less research exists on the empirical question of how movements in the natural rate affect macroeconomic indicators in the short and long runs. Does a fall in the natural rate increase or decrease the trend of output? Is a fall in the natural rate contractionary or expansionary? Is it inflationary or deflationary? This paper aims to address these questions.

We contribute to the related empirical literature by allowing the permanent component of the real interest rate to be a source of movements in real activity and prices. This approach is motivated by the dynamic optimizing general equilibrium literature in which low frequency movements in the real interest rate can occur due to exogenous permanent movements in subjective discount rates, demographic trends, or other exogenous factors determining secular shifts in the propensities to save of domestic or foreign agents.

With this motivation in mind, we develop an empirical model that combines elements of a structural vector autoregression (SVAR) model and a dynamic stochastic general equilibrium (DSGE) model. Like standard SVAR formulations, the model contains a relatively small number of cross-equation restrictions. And like DSGE structures, the model allows for identified trend shocks—including permanent disturbances to the natural rate—to affect not only the trends of endogenous variables but also their cyclical components. These features make it possible to estimate jointly the long-run and business-cycle effects of movements in the permanent component of the real interest rate. Further, the proposed formulation accommodates more identified shocks than observables. This is important because it creates a competition for the data between the permanent natural rate shock and a rich set of alternative disturbances with a precise economic interpretation.

The model incorporates three permanent shocks: a permanent real interest rate shock, a permanent productivity shock, and a permanent monetary shock. In addition, the model includes a transitory monetary shock and a transitory real shock. We estimate the model using Bayesian techniques on annual U.S. data for output, inflation, and the short-term nominal interest rate over the period 1900 to 2021.

The main finding of the paper is that a shock that permanently lowers the real interest rate permanently lowers the trend path of output and is contractionary and deflationary already in the short run. Specifically, a one-percentage-point fall in the permanent component of the real interest rate reduces the trend level of output by about 8 percentage points. Of this 8-percentage-point fall in the trend of output two-thirds are estimated to take place in the short run (within a year). In addition, inflation falls on impact by 25 basis points and remains below target for about 3 years. These findings suggest that the identified natural rate shock is a demand side shock as it moves output and inflation in the same direction.

Importantly, we estimate that negative natural rate shocks have a negative effect on output trend even when the economy is not constrained by the zero lower bound on policy rates. This finding suggests that the mechanism that triggers the effect may be more general than the one invoked by the secular stagnation hypothesis in its different varieties. This is so because in this theory being near the ZLB is key for a fall in the natural rate to push the economy onto a lower trend trajectory.

We further find that the permanent component of the real interest rate is driven only by real factors. An implication of this finding is that the Fisher effect holds, that is, monetary policy does not affect the real interest rate in the long run. Applied to the monetary policy actions observed in the United States and the euro area at the time of this writing, this result says that a normalization of nominal interest rates need not mean the end of low real rates.

The main results of the paper are shown to be robust to a number of modifications. In particular, we estimate the model on postwar data, allow for heteroskedasticity in measurement error, and consider alternative identification schemes.

The present investigation is related to two strands of literature. One strand is concerned with the estimation of the time path of the natural rate of interest. A seminal contribution in this body of work is Laubach and Williams (2003), who model the natural rate as an unobserved permanent component in the real interest rate. Laubach and Williams (2016) apply this model and estimate that the natural rate in the United States has experienced a persistent decline since the 1980s with a particularly large fall around the financial crisis of 2008. Holston et al. (2017), Del Negro et al. (2019), Ferreira and Shousha (2021), Cesa-Bianchi et al. (2022), and Hamilton et al. (2016) provide international evidence on the persistent decline of the natural rate over the few decades. The latter paper estimates a larger degree of uncertainty around the level of the natural rate relative to other papers in the literature. Del Negro et al. (2017) find that the decline in the natural rate in the United States was primarily caused by changes in the convenience yield of safe assets. The contribution of the present paper to this strand of the literature is to estimate the effects of innovations in the natural rate of interest (i.e., innovations to the permanent component of the real interest rate) on output, its trend, and inflation. To our knowledge this represents the first attempt to characterize empirically the macroeconomic effects of permanent real interest rate shocks.

The other strand of literature to which this paper is related is theoretical and goes back to the contributions of Hansen (1939) who linked secular declines in the real interest rate to permanent declines in potential output when the economy is near or at the zero lower bound on interest rates. Modern formulations of this hypothesis have been advanced by Summers (2014), Eggertsson et al. (2019), and Benigno and Fornaro (2018) among others. In all of these formulations, the zero lower bound is a key ingredient in the formation of a secular stagnation. The contribution of the present paper in this regard is the finding that the secular stagnation phenomenon appears to be more general, in the sense that negative natural rate shocks appear to depress the trend path of output even when the economy is reasonably far away from the ZLB. Finally, the econometric framework extends the one developed in Uribe (2022) to allow for shocks to the permanent component of the real interest rate.

The remainder of the paper is organized as follows. Section 2 presents the empirical model and the identification strategy. It also describes the data and the estimation procedure. Section 3 presents the main results of the paper, and section 4 contains a number of robustness checks. Section 5 concludes.

# 2 Empirical Model, Identification, and Estimation

In this section, we present the semi-structural empirical model and discuss the identification scheme, the data, and the estimation strategy.

## 2.1 Empirical Model

The model structure is based on Uribe (2022). It departs from that framework by incorporating permanent disturbances to the real interest rate, which are the focus of the present analysis. The model is cast in terms of the logarithm of real output per capita, denoted $y_t$, the inflation rate, denoted $\pi_t$, and the short-term nominal interest rate, denoted $i_t$. For simplicity, the exposition of the model omits intercepts. We suppose that output is cointegrated with a nonstationary real shock $X_t$ and a permanent real-interest-rate shock $X_t^r$. Specifically, we define the unobservable stationary variable $\hat{y}_t$ as

$$\hat{y}_t = y_t - X_t - \delta X_t^r.$$

The endogenous latent variable $\hat{y}_t$ is interpreted as the cyclical component of output, the exogenous latent variable $X_t$ captures permanent movements in the state of technology, and the exogenous latent variable $X_t^r$ captures permanent changes in the natural rate of interest originating from non-monetary forces. Permanent movements in the natural rate of interest due to changes in $X_t^r$ could stem from, for example, secular variations in demographic variables, exogenous changes in subjective discount rates, or in other factors determining the domestic or external willingness to save. The output trend is then given by

$$\text{output trend} = X_t + \delta X_t^r + \text{constant} \times t. \tag{1}$$

The parameter $\delta$ governs the long-run effect of permanent real-interest-rate shocks on the trend of output. As will become clear shortly, a fall in $X_t^r$ corresponds to a decline in the natural rate of interest. Thus, a positive value of $\delta$ implies that a permanent decline in the natural rate of interest lowers the trend of output.

The model assumes that inflation is cointegrated with an exogenous stochastic nonstationary nominal disturbance denoted $X_t^m$. Specifically, we define the cyclical component of inflation, $\hat{\pi}_t$, as

$$\hat{\pi}_t = \pi_t - X_t^m.$$

The variable $\hat{\pi}_t$ is stationary. The permanent monetary shock $X_t^m$ can be interpreted as the permanent component of a stochastic inflation target.

The model assumes that the nominal interest rate is cointegrated with the inflation target $X_t^m$ and with the permanent real-interest-rate shock $X_t^r$, so that its cyclical component, denoted $\hat{i}_t$, is given by

$$\hat{i}_t = i_t - (1+\alpha)X_t^m - X_t^r.$$

The parameter $\alpha$ determines the long-run effect of permanent changes in the inflation target on the real interest rate. When $\alpha$ equals zero, long-run neutrality holds in the sense of Fisher (1896). In this case, a normalization of nominal rates, like the one seemingly underway in much of the developed world at the time of this writing, would have no permanent effect on the real interest rate. A nonzero value of $\alpha$ introduces a monetary permanent component into the real interest rate. For example, a positive value of $\alpha$ would imply that the persistent disinflation observed in the United States since the Volcker era and up until the beginning of the COVID-19 pandemic was in part responsible for the concurrent fall in real interest rates.

Defining the natural rate of interest as the real interest rate prevailing in a stochastic steady state—that is, along a path in which the cyclical components of output, inflation, and the interest rate are all nil—we have that

$$\text{natural rate of interest} = X_t^r + \alpha X_t^m. \tag{2}$$

It is then clear that $X_t^r$ represents the permanent component of the real interest rate stemming from nonmonetary factors. When $\alpha = 0$, i.e., when the long-run Fisher effect holds, $X_t^r$ is itself the permanent component of the real interest rate.[1]

In addition to the three permanent shocks, $X_t$, $X_t^m$, and $X_t^r$, the model includes a stationary monetary shock, denoted $z_t^m$, and a stationary real shock, denoted $z_t$. The dynamics of the cyclical components of the three endogenous variables of the model are assumed to be given by the following first-order autoregressive system

$$\begin{bmatrix} \hat{y}_t \\ \hat{\pi}_t \\ \hat{i}_t \end{bmatrix} = B \begin{bmatrix} \hat{y}_{t-1} \\ \hat{\pi}_{t-1} \\ \hat{i}_{t-1} \end{bmatrix} + C \begin{bmatrix} \Delta X_t^m \\ z_t^m \\ \Delta X_t \\ z_t \\ \Delta X_t^r \end{bmatrix}. \tag{3}$$

This structure is inspired by that of optimizing dynamic models, in which the equilibrium values of the cyclical components of endogenous variables depend not only on realizations of stationary shocks but also on disturbances to trend components. In particular, this feature of the model allows us to estimate the impulse responses of output, inflation, and the interest rate to an innovation in $X_t^r$. It is also an important difference with VAR models in which trend shocks are not allowed to affect the cycle. (See, for example, Del Negro et al., 2017. In terms of the present notation, that paper assumes that the first, third, and fifth columns of $C$ are nil.)

The stochastic driving vector $\begin{bmatrix} \Delta X_t^m & z_t^m & \Delta X_t & z_t & \Delta X_t^r \end{bmatrix}'$ follows the first-order autoregressive law of motion

$$\begin{bmatrix} \Delta X_{t+1}^m \\ z_{t+1}^m \\ \Delta X_{t+1} \\ z_{t+1} \\ \Delta X_{t+1}^r \end{bmatrix} = \rho \begin{bmatrix} \Delta X_t^m \\ z_t^m \\ \Delta X_t \\ z_t \\ \Delta X_t^r \end{bmatrix} + \Psi \begin{bmatrix} \epsilon_{t+1}^{X^m} \\ \epsilon_{t+1}^{z^m} \\ \epsilon_{t+1}^{X} \\ \epsilon_{t+1}^{z} \\ \epsilon_{t+1}^{X^r} \end{bmatrix}, \tag{4}$$

where $\rho$ and $\Psi$ are diagonal matrices and $\epsilon_t^s$, for $s = X^m, z^m, X, z, X^r$, are i.i.d. disturbances distributed $N(0,1)$. This component of the model is also inspired by DSGE formulations in which all exogenous shocks are identified and follow distinct laws of motion (more on identification in subsection 2.2).

The model is cast in terms of latent variables. To be able to estimate it—that is, to

[1]In this case, $X_t^r$ corresponds to the variable $r_t^*$ in the notation of related empirical studies (e.g., Laubach and Williams, 2003; Del Negro et al., 2017).

estimate the matrices $B$, $C$, $\rho$, $\Psi$ and the parameters $\alpha$ and $\delta$—we introduce a vector of observable variables, for which the model has precise predictions. Specifically, we assume that the econometrician observes with measurement error the growth rate of real GDP per capita, denoted $\Delta y_t$, the change in the inflation rate, denoted $\Delta\pi_t$, and the change in the short-term nominal interest rate, denoted $\Delta i_t$. The observation equations are given by the following three identities:

$$\Delta y_t = \hat{y}_t - \hat{y}_{t-1} + \Delta X_t + \delta \Delta X_t^r + \mu_t^y, \tag{5}$$

$$\Delta \pi_t = \hat{\pi}_t - \hat{\pi}_{t-1} + \Delta X_t^m + \mu_t^\pi, \tag{6}$$

and

$$\Delta i_t = \hat{i}_t - \hat{i}_{t-1} + (1+\alpha)\Delta X_t^m + \Delta X_t^r + \mu_t^i, \tag{7}$$

where the vector $[\mu_t^y \; \mu_t^\pi \; \mu_t^i]'$ is a normally distributed mean-zero i.i.d. measurement error with a diagonal variance-covariance matrix denoted $R$. Observation equations (5)–(7) say that, up to a measurement error, the growth rate of an observed variable can be expressed as the sum of the growth rate of its cyclical component and the growth rate of its trend component.

## 2.2 Identification and Priors

The cointegration restrictions discussed thus far are themselves identification restrictions for the nonstationary shocks. In particular, they imply that $X_t$ is the only shock that has a long-run effect on output but not on inflation or the nominal interest rate; $X_t^m$ is the only shock that can have a long-run effect on inflation; and $X_t^r$ is the only shock that can have a long-run effect on output and the nominal interest rate but not on inflation.

To identify the stationary shocks $z_t^m$ and $z_t$, we proceed as follows: We assume that the stationary monetary shock $z_t^m$ does not affect output or inflation on impact, that is, we impose

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

in equation (3). Without loss of generality, we normalize the impact effects of the stationary monetary shock on the interest rate and of the stationary real shock ($z_t$) on output to unity, that is, we set

$$C_{14} = C_{32} = 1.$$

The assumption that $z_t^m$ has a zero impact effect on output and inflation may sound too restrictive, given the annual frequency of the data. However, in quarterly data, the effects of monetary shocks often manifest with a delay of several quarters (see, for example, Christiano et al., 2005). As a robustness check, in section 4.4 we consider an alternative identification scheme for $z_t^m$ in which stationary monetary contractions are allowed to have a nonpositive impact effect on output and inflation.

We assume that at the mean of the prior distribution the long-run Fisher effect holds, that is, $\alpha = 0$ in observation equation (7). Specifically, we assume that $\alpha$ has a normal prior distribution with mean 0 and standard deviation 0.25. This means that if $\alpha$ is one standard deviation above its prior mean, then a permanent increase in inflation of 1 percentage point leads to a permanent increase in the nominal interest rate of 1.25 percentage points and therefore to a permanent increase in the real interest rate of 0.25 percentage points.

The assumption that the prior mean of $\alpha$ is zero is motivated by the following three observations. First, across time and countries over long periods of time, inflation and the nominal interest rate tend to move one for one (see, for example, Uribe, 2022, Figures 1 and 2). Second, existing cointegration analyses are inconclusive about the sign of the deviation from the long-run Fisher effect, with some finding that a permanent increase in the inflation rate leads in the long run to a permanent increase in the real interest rate (Azevedo et al., forthcoming), to a decrease (King and Watson, 1997), or to no change (Uribe, 2022). Third, a large number of studies on the joint behavior of output, inflation, and the nominal interest rate assumes, on theoretical grounds, that the real interest rate is independent of the rate of inflation in the long run (see, for example, Galí, 1992, among many others).

We impose a zero prior mean for the parameter $\delta$ in observation equation (5), which implies a prior belief that the real interest rate shock $X_t^r$ has no effect on the trend of output. Specifically, we assume that $\delta$ has a normal prior distribution with mean 0. In addition, we assume that this distribution is quite diffuse by setting a standard deviation of 5, which implies that if $\delta$ is one standard deviation above its prior mean, then an $X_t^r$ shock that lowers the permanent component of the real interest rate by 1 percentage point reduces the trend of output by 5 percent. The rationale behind adopting a diffuse prior for $\delta$ is a body of theoretical work, going back to Hansen (1939), that attributes large variations in potential output to movements in the natural rate of interest. The rationale for the assumed symmetry around zero is that, while models of secular stagnation are designed to deliver a negative effect of a drop in the natural rate on trend output, it is easy to show that in the neoclassical growth model this effect is positive and, under plausible calibrations, also large.

The remaining prior assumptions follow Uribe (2022). The elements of the matrix $B$ in equation (3) are assumed to have normal prior distributions. In the spirit of the Minnesota prior, we impose a prior mean of 0.5 on the diagonal elements of $B$ and a prior mean of 0 on the off diagonal elements. All elements of $B$ are assumed to have a prior standard deviation of 0.3.

Next, we present the assumed prior distributions of the estimated elements of the matrix $C$ in equation (3). Under the mean of the prior distribution, an innovation to $X_t^m$ is assumed to have a zero impact effect on inflation and the interest rate. This requires setting the prior means of $C_{21}$ and $C_{31}$ to -1. All other estimated elements of $C$ are assumed to have a prior mean of 0. Further, the estimated elements of $C$ are assumed to have normal prior distributions with standard deviations equal to 1.

Consider now the prior distributions of the estimated parameters of the laws of motion of the exogenous driving forces given in equation (4). The diagonal elements of the matrix $\rho$ are assumed to have a beta prior distribution with mean 0.3 and a standard deviation of 0.2, with the exception of element $(4, 4)$, which is assumed to have a mean of 0.5 and a standard deviation of 0.2. The rationale behind the assumed mean values of $\rho_{11}, \rho_{33}$, and $\rho_{55}$ is that changes in output, inflation, and the nominal interest rate have relatively low serial correlations. The rationale for the prior mean of $\rho_{22}$ is that transitory monetary disturbances are typically assumed to be i.i.d. or to have a low persistence. (For example, Smets and Wouters, 2007, assign this parameter a prior mean of 0.5 for a model estimated on quarterly data, which at an annual frequency corresponds to a value of about 0.2.) The justification for the higher prior mean for the serial correlation of $z_t$ is that in business cycle analysis stationary productivity shocks are typically estimated to have a relatively high persistence. The standard deviations of all exogenous shocks, given by the diagonal of the matrix $\Psi$, are assumed to have gamma prior distributions with means and standard deviations equal to 1.

The trend of output is assumed to have a deterministic component. The growth rate of this component, given by the intercept of equation (5) (not shown), is assumed to have a normal prior distribution with a mean and a standard deviation equal to the sample mean and standard deviation of output growth. Changes in inflation and the interest rate are assumed to have a zero mean. That is, observation equations (6) and (7) are assumed to have no intercept. Finally, the diagonal elements of the variance-covariance matrix $R$ of measurement errors in observation equations (5)–(7) are assumed to have uniform prior distributions. The lower bounds of these distributions are set to zero, and the upper bounds are set to ensure that under the prior distribution measurement errors account for no more than 10 percent of the variance of the data.

Table 1 provides a summary of the assumed prior distributions.

## 2.3 Data and Estimation

The model is estimated on annual U.S. data from 1900 to 2021. For the period 1900 to 2017 the data is taken from the Jordá-Schularick-Taylor Macrohistory Database, see Jordá

Table 1: Prior Distributions

| Parameter | Distribution | Mean | Std. Dev. |
|---|---|---|---|
| Diagonal elements of $B$ | Normal | 0.5 | 0.3 |
| Off diagonal elements of $B$ | Normal | 0 | 0.3 |
| $C_{21}$, $C_{31}$ | Normal | -1 | 1 |
| All other estimated elements of $C$ | Normal | 0 | 1 |
| $\rho_{ii}$, $i = 1, 2, 3, 5$ | Beta | 0.3 | 0.2 |
| $\rho_{44}$ | Beta | 0.5 | 0.2 |
| Diagonal elements of $\psi$ | Gamma | 1 | 1 |
| $\alpha$ | Normal | 0 | 0.25 |
| $\delta$ | Normal | 0 | 5 |
| Diagonal elements of $R$ | Uniform | $\frac{\text{var}(o_t)}{10\times 2}$ | $\frac{\text{var}(o_t)}{10\times\sqrt{12}}$ |
| Estimated element of $A$ | Normal | $E(\Delta y_t)$ | $\sqrt{\frac{\text{var}(\Delta y_t)}{T}}$ |

Notes. $T$ denotes the sample length. The vector $o_t$ contains the observables, $o_t = [\Delta y_t \ \Delta\pi_t \ \Delta i_t]'$. The vector $A$ denotes the mean of the vector of observables, $A = E(o_t)$.

et al. (2017). Specifically, we downloaded from that database the series rgdppc, cpi, and stir corresponding, respectively, to real GDP per capita, the consumer price index, and the nominal short-term interest rate. Since 1955 the short-term nominal interest rate measure corresponds to the effective federal funds rate. For the period 2018 to 2021, the data source for real GDP per capita is the Bureau of Economic Analysis, for the consumer price index the Bureau of Labor Statistics, and for the federal funds rate the Federal Reserve Board. The annual value of the consumer price index is computed as the arithmetic average over the corresponding monthly observations. Output growth, inflation, and the short-term interest rate are expressed in percent per year.

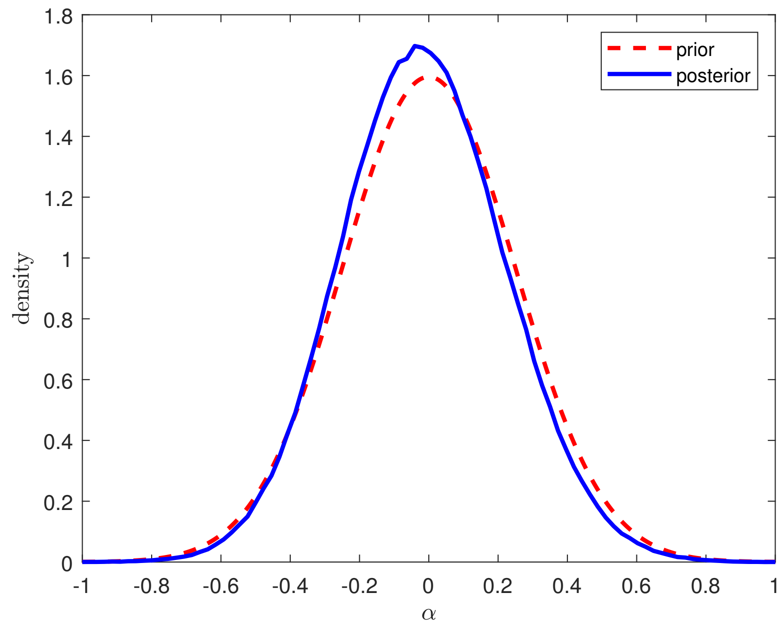
We estimate the model using Bayesian methods and use the random walk Metropolis-Hastings algorithm to obtain 50 million draws from the posterior distribution of the estimated parameter vector. We target an acceptance rate between 0.2 and 0.4. The results presented in the remainder of the paper are based on random subsamples from the 50-million draws of length either 1 million or 100 thousand, depending on the computation time involved in the quantitative analysis.

# 3 Results

This section addresses the question of how innovations in the permanent component of the real interest rate affect output and inflation in the long and short runs. Before plunging into these issues, it is necessary to establish whether the permanent component of the real interest rate is affected by monetary factors.

## 3.1 Does Monetary Policy Affect the Natural Rate of Interest?

Figure 1: Prior and Posterior Densities of $\alpha$




Notes. The parameter $\alpha$ measures the effect of an increase in trend inflation on the natural rate of interest, see equation (2). A value of zero implies that the long-run Fisher effect holds.

The Great Moderation, which lasted from the mid 1980s until the global financial crisis of 2008, was characterized by a significant fall in both the inflation rate and the real interest rate. This phenomenon raises the question of whether there is a connection between the permanent component of the inflation rate, $X_t^m$, and the natural rate of interest, $X_t^r + \alpha X_t^m$. Because $X_t^r$ and $X_t^m$ are independent from each other, the question of whether the permanent component of inflation affects the natural rate of interest reduces to ascertaining whether the parameter $\alpha$ is significantly different from zero.

Figure 1 displays the prior and posterior distributions of $\alpha$. The two distributions are centered around zero and show a significant amount of overlap. We interpret this result as suggesting that the posterior distribution does not contradict the prior assumption that $\alpha$ has a mean value of zero. From an economic point of view this result suggests that it is reasonable to maintain that the long-run Fisher effect holds, that is, that a change in the permanent component of inflation has no lasting effect on the real interest rate. This result should not be surprising in light of the discussion of section 2.2 documenting that the existing empirical literature has had a hard time establishing a consistent long-run connection between monetary policy and real interest rates.

It is then reasonable to maintain that the natural rate of interest is driven solely by real factors embodied in the exogenous variable $X_t^r$. Under this interpretation, the concomitant fall in inflation and real interest rates observed during the Great Moderation period was a coincidence. This view is not at odds with conventional narratives: On the monetary side during the Great Moderation the Federal Reserve consolidated the Volcker disinflation. This policy was arguably motivated, not by real factors, but by a desire to re-anchor inflationary expectations gone adrift in the 1970s. On the real side, a number of developments, including an aging population, the emergence of China as a major supplier of funds to world financial markets, and changes in the distribution of wealth within the United States, are often cited as factors that affected the level of the natural rate of interest.

The finding that $\alpha$ is not significantly different from zero has policy implications. It implies, for example, that normalization of policy rates from exceptionally low levels to higher values closer to historical averages, like the ones seemingly underway in the United States and other advanced economies at the time of this writing, per se should not be expected to mean the end of low real rates.

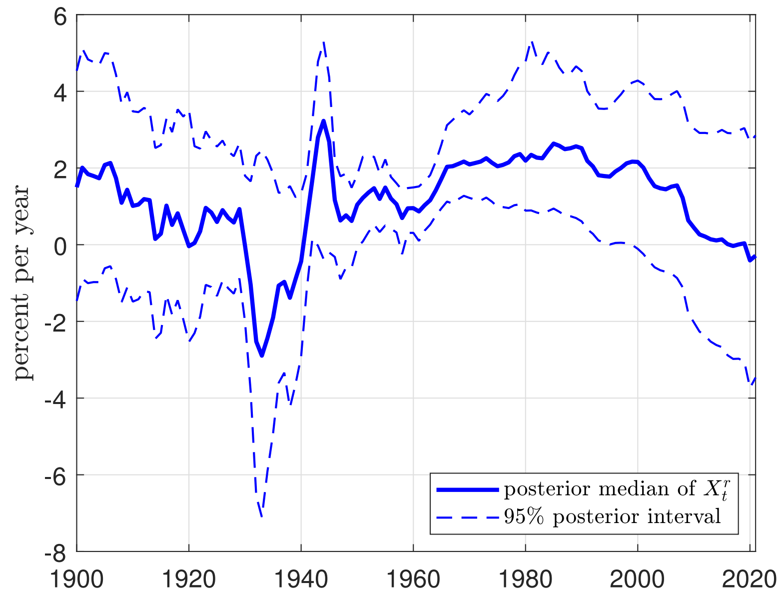
In light of the result that monetary policy has no significant long-run effects on the real interest rate, in what follows we will refer to the real permanent component of the real interest rate, $X_t^r$, as the natural rate of interest.

## 3.2 Natural Rate Supercycles

A large number of papers devoted to the estimation of the natural rate of interest have documented a persistent decline in this variable over the past decades. Figure 2 offers a long run perspective on the behavior of the natural rate, $X_t^r$. It displays its inferred path over the period 1900-2021. The variable $X_t^r$ is computed by Kalman smoothing using 100,000 draws from the posterior chain of parameter vectors of length 50 million. This method delivers an estimate of $X_t^r$ up to a constant. To facilitate interpretation, we set this constant to ensure that $X_t^r$ and the observed ex-post real rate, $i_t - \pi_{t+1}$, have the same sample mean, 1.15 percent per year. The solid line plots the posterior median of $X_t^r$ and the broken lines indicate the 2.5th and 97.5th percentiles of its posterior estimate.

The figure shows that the natural rate of interest displays long cycles. We refer to those as natural rate supercycles. For the estimation period, the first supercycle begins in 1900 and ends in the mid 1980s. The trough of this supercycle coincides with the trough of the

Figure 2: The Natural Rate of Interest: 1900-2021

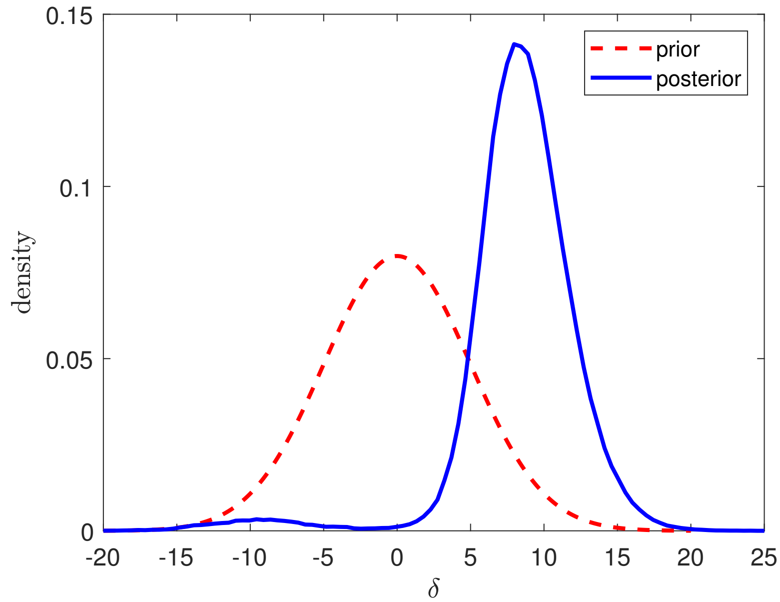



Notes. The variable $X_t^r$ is computed by Kalman smoothing using 100,000 draws from the posterior chain. It is normalized by adding a constant to match the 1.15 percent per year sample mean of $i_t - \pi_{t+1}$. The solid line is the median of $X_t^r$ and the broken lines indicate the 2.5th and 97.5th percentile of $X_t^r$, respectively.

Great Depression in the 1930s. The figure suggests that the fall in the natural rate that has been taking place over the past three decades is part of a second supercycle that began in the mid 1980s.

The long swings in the natural rate of interest predicted by the model deliver two important insights. First, there were significant movements in the natural rate of interest pre 1980. This finding departs from that stressed in Del Negro et al. (2019). These authors estimate, also using long data but a different empirical strategy, that the natural rate of interest displayed little variation prior to the Great Moderation. One possible reason for the difference in results is that these authors impose a tight prior distribution for the variance of the innovation to the trend component with a small mean. This assumption about the prior distribution punishes movements in the trend component of the real interest rate. Second, the persistent fall in the natural rate of interest over the past three decades is not unprecedented. Between 1900 and 1933, the natural rate fell by 4 percentage points. This figure is larger than the estimated 3 percentage points decline in the natural rate since the beginning of the most recent downward swing of the supercycle in 1985.

Figure 3: Prior and Posterior Densities of $\delta$



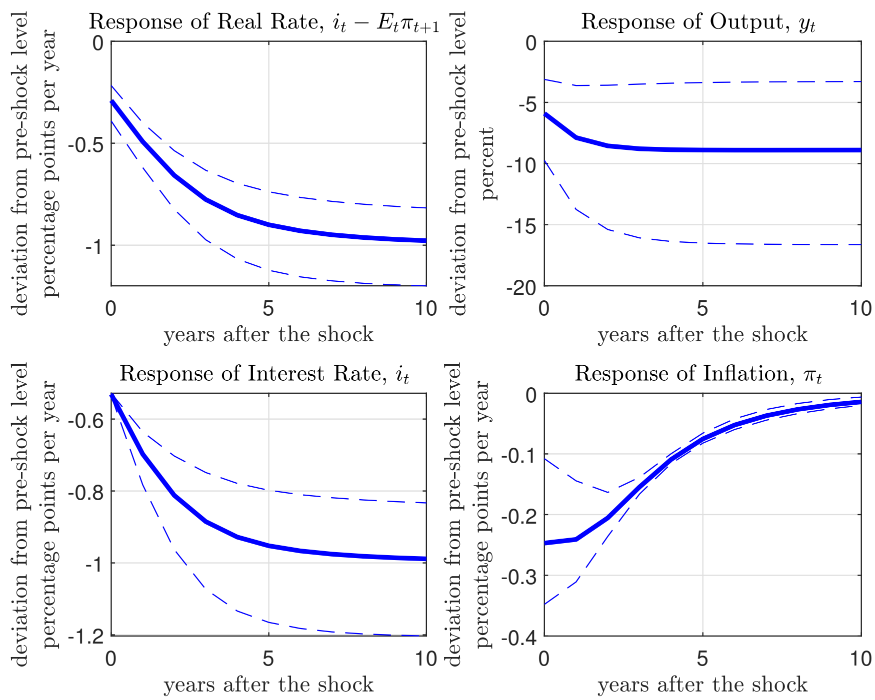

Notes. The parameter $\delta$ measures the effect of a permanent change in the natural rate of interest, $X_t^r$, on the trend of output. A positive value of $\delta$ means that a decline in the natural rate of interest (a fall in $X_t^r$) lowers the trend of output.

## 3.3 Effects of Natural Rate Shocks on the Trend of Output

This subsection presents the main results of the paper. According to equation (1), the effect of changes in the permanent component of the real interest rate, $X_t^r$, on the trend of output, $X_t + \delta X_t^r$, is governed by the parameter $\delta$. Figure 3 displays the prior and posterior distributions of this parameter. The data appears to favor positive values of $\delta$. Under the posterior distribution, the probability that $\delta$ is positive is 97 percent compared to 50 percent under the prior distribution. The posterior median of $\delta$ is 8.6, which means that a one percentage point fall in the natural rate of interest $X_t^r$ lowers the trend level of output by 8.6 percent. This finding suggests a sizable degree of confidence that a fall in the natural rate has a negative long-run effect on the level of output per capita.

We note, however, that the posterior distribution of $\delta$ does have some mass (about 3 percent) to the left of zero, and that in this range it displays a second mode. The long left tail can give rise to wide confidence bands of objects such as impulse response functions. For this reason, in the dynamic analysis that follows we will focus on model predictions conditional on positive values of $\delta$.

Figure 4: Impulse Response to a Decline in the Natural Rate of Interest



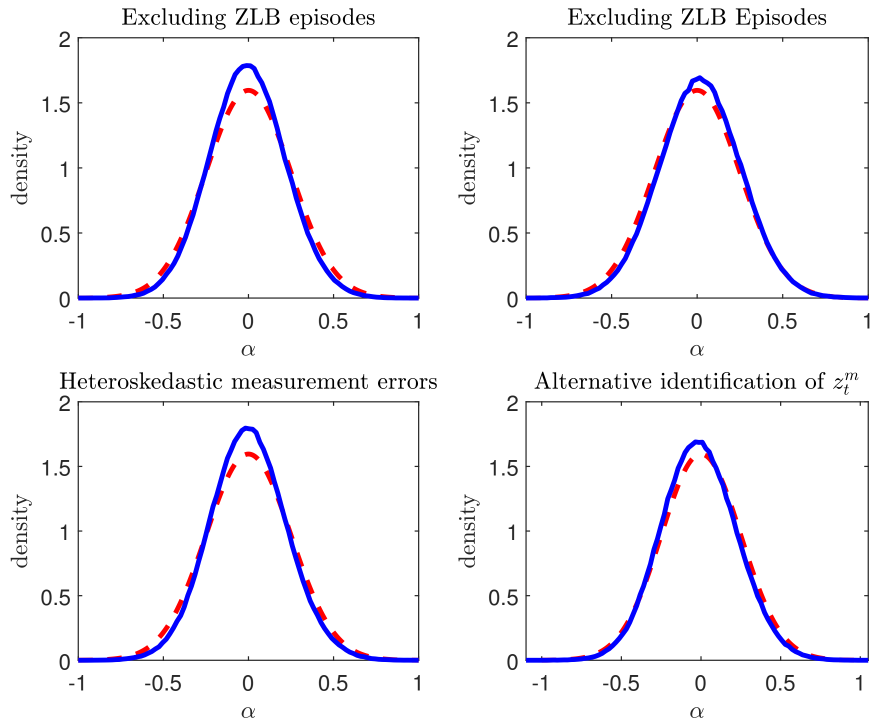

Notes. Solid lines display the posterior mean response to a negative natural rate shock (a decrease in $X_t^r$) that lowers the real interest rate by 1 annual percentage point in the long run. Broken lines are asymmetric 95-percent confidence bands computed using the Sims-Zha (1999) method from 100,000 draws from the posterior distribution. Impulse responses and confidence bands are conditional on $\delta > 0$.

## 3.4 Short-Run Effects of Natural Rate Shocks

What does the transition to a lower natural rate look like? Figure 4 addresses this question. It displays the posterior mean response to a negative natural rate shock (a fall in $X_t^r$) that lowers the real interest rate by 1 percentage point in the long run. The figure also includes asymmetric 95-percent confidence bands computed using the Sims-Zha (1999) method. Per the discussion in subsection 3.3, the plotted impulse responses and their associated error bands are conditional on $\delta > 0$. The figure is largely unchanged when one does not condition on positive $\delta$ values except that the error band around the output impulse response is significantly wider.

The top right panel of the figure shows that a fall in the permanent component of the real interest rate is contractionary in the short run. On impact output falls by 5.6 percent, which represents two thirds of its long run decline of 8.3 percent (the posterior mean of $\delta$). The fall in the natural rate is also deflationary in the short run (bottom right panel of Figure 4). A one percentage point fall in the natural rate lowers inflation by 25 basis

Figure 5: Prior and Posterior Densities of $\alpha$: Sensitivity Analysis




Notes. Solid lines display the posterior density of the parameter $\alpha$ and broken lines the prior density.

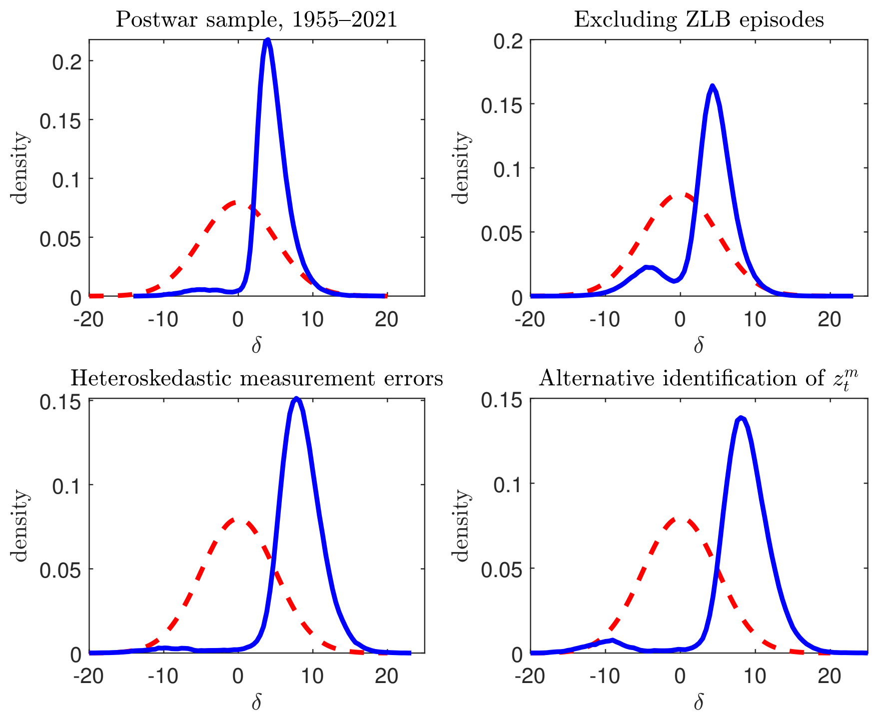
points on impact. The decline in inflation is persistent with a half life of about 4 years. Throughout the transition, the nominal rate falls by more than inflation, implying that a permanent decline in the natural rate of interest leads to a reduction in the real interest rate, $i_t - E_t\pi_{t+1}$, not only in the long run but also in the short run (left panels of Figure 4).

Taken together, the findings presented in this section suggest that a fall in the natural rate of interest causes a downward parallel shift in the trend of the logarithm of output and a contraction and deflation in the short run.

# 4 Robustness Analysis

This section presents a number of exercises aimed at ascertaining the robustness of the main results to estimating the model over the postwar sample, excluding periods with zero lower bound episodes, allowing for heteroskedasticity in measurement errors, and imposing alternative identification schemes monetary shocks. The results are presented in Figures 5 through 7.

Figure 6: Prior and Posterior Densities of $\delta$: Sensitivity Analysis



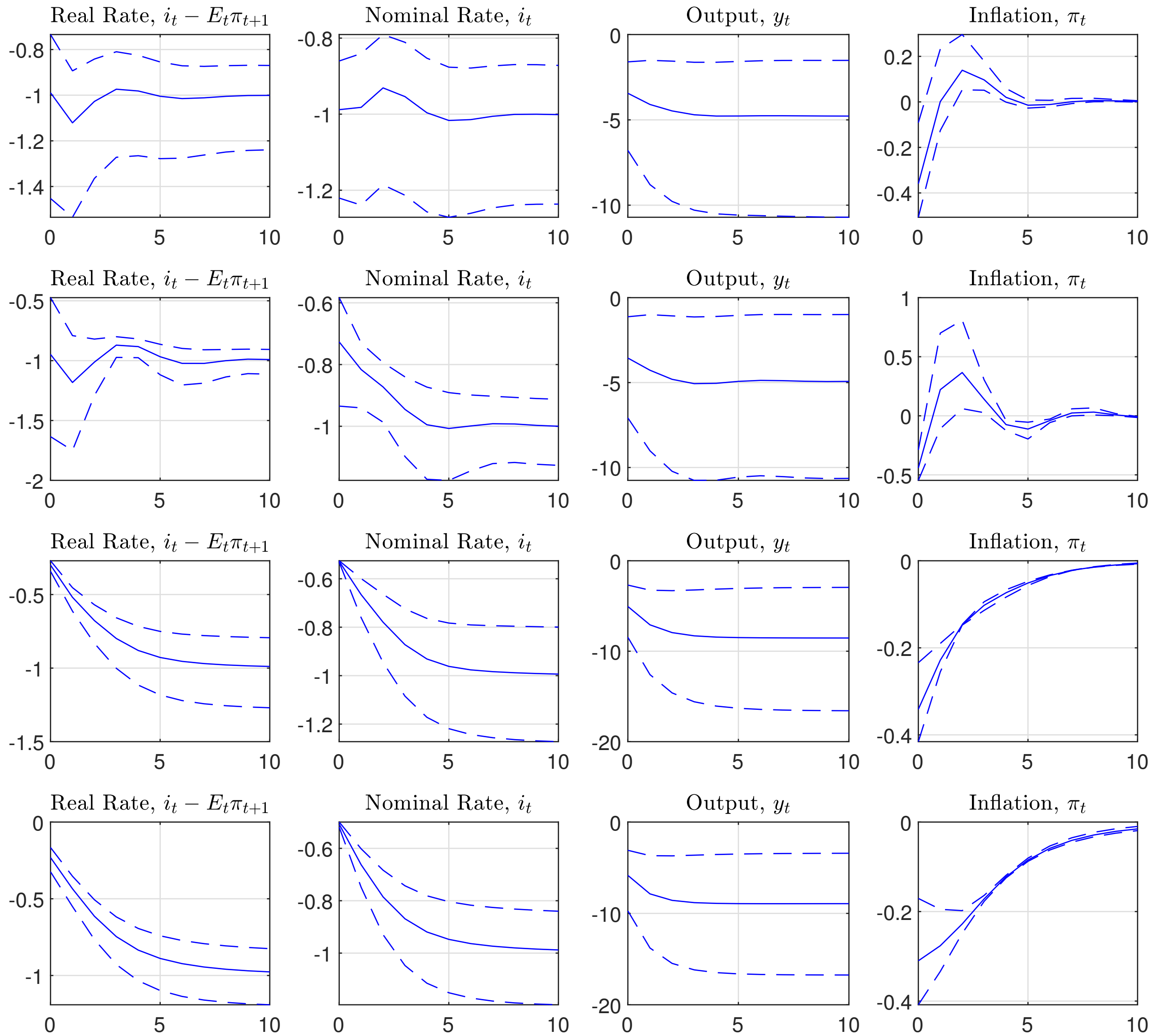

Notes. Solid lines display the posterior density of the parameter $\delta$ and broken lines the prior density.

## 4.1 Postwar Sample: 1955 to 2021

We begin the sensitivity analysis by estimating the model using a shorter sample starting in 1955 and ending in 2021. There are pros and cons involved in excluding the first half of the sample. On the positive side, it addresses concerns about changes in the structure of the economy, monetary policy, and the persistence and volatility of aggregate disturbances since 1900. An additional advantage of focusing on the postwar sample period is that it is more commonly used in business-cycle analysis. On the negative side, significantly shortening the sample hinders the estimation of permanent disturbances. It also makes it harder to detect whether the fall in the natural rate observed in the past two decades is a unique or recurrent phenomenon.

The top-left panels of Figures 5 and 6 and the first row of Figure 7 show that the main results of the paper hold when the model is estimated using postwar data. In particular, the fact that $\alpha$ continues to be insignificantly different from zero suggests that the Fisher effect holds in the long run. The fact that the estimate of the parameter $\delta$ is significantly greater than 0 implies that a fall in the natural rate has a significant negative effect on the trend of output. The posterior median of $\delta$, however, is lower than its long-data counterpart (4.4 versus 8.6). The impulse response functions to a negative $X_t^r$ shock that lowers the real rate

Figure 7: Impulse Responses to a Decline in the Natural Rate: Sensitivity Analysis




Notes. Row 1, postwar sample (1955 to 2021); row 2, excluding ZLB episodes; row 3, heteroskedastic measurement errors; and row 4, alternative identification of the stationary monetary shock $z_t^m$. The fall in $X_t^r$ is such that the real rate falls by 1 percentage point in the long run. Solid lines are posterior means, and broken lines are asymmetric 95-percent confidence bands computed with the Sims-Zha (1999) method and using 100,000 draws from the posterior distribution conditional on $\delta > 0$.

by 1 percentage point in the long run show that the short-run effect of a fall in the natural rate continues to be deflationary and contractionary.

## 4.2 Excluding ZLB Episodes

The secular stagnation hypothesis argues that when the policy rate is near its effective lower bound, in response to negative natural rate shocks the real interest rate may not be able to fall to the level required for aggregate demand to meet aggregate supply. As a consequence, in such circumstances, the economy suffers involuntary unemployment. A natural question is whether the ZLB is a necessary condition for negative natural rate shocks to negatively affect the trend path of output. Put differently, can negative natural rate shocks push the economy into a secular stagnation even when the policy rate is not near its effective lower bound? In the United States, an example of a continuous period in which the policy rate was away from its effective lower bound runs from the end of World War II until the eve of the global financial crisis. Accordingly, we reestimate the model using data from 1946 to 2007.

The top right panel of Figure 6 displays the resulting posterior density of $\delta$. The shape of this distribution is bimodal with a peak at a positive value of 4.3 and another at a negative value of -4.6. However, the bulk of the posterior probability of $\delta$, 85 percent, lies in the positive range, and the posterior median of $\delta$ coincides with the positive peak of 4.3. This finding suggests that even when the economy is away from the zero lower bound, there is a substantial probability that a permanent fall in the real interest rate will cause a fall in the trend of output.

The short-run effects of natural rate shocks are also robust to estimating the model on a sample in which the policy rate is not constrained by the zero lower bound. The second row of Figure 7 presents the impulse responses of output, inflation, and nominal and real interest rates to an $X_t^r$ shock that lowers the real interest rate by 1 percentage point in the long run conditional on $\delta$ being positive. The impulse responses suggest that negative natural rate shocks continue to be contractionary in the short run.

Overall, the message conveyed by this robustness check is that the ZLB, which is a key element of the secular stagnation hypothesis, does not appear to be a necessary condition for negative natural rate shocks to cause a downward shift in the trend path of output or a recession. This finding suggests that more theory is required to understand the macroeconomic effect of shocks to the natural rate of interest. In particular, the results point to a need for a theory that explains why permanent declines in the real interest rate lower the permanent component of output and cause recession in the short run, even when the policy rate is away from its effective lower bound.

## 4.3 Heteroskedasticity in Measurement Errors

The third robustness check we perform is to allow for heteroskedasticity in measurement errors pre and post 1955. The rationale for this exercise is that arguably, the systematic compilation of NIPA data and aggregate price indices began in earnest after World War II, which conceivably could have resulted in larger measurement errors in the earlier period. Accordingly, we estimate the model allowing measurement errors to capture up to 20 percent of the variance of each observable prior to 1955. This is twice the amount allowed in the baseline estimation. In the post 1955 subperiod, we continue to limit the fraction of the variance of the observables explained by measurement error at 10 percent.

We find that measurement errors are indeed much larger in the pre-postwar period. The fraction of the variance of the observables explained by measurement error is 10 times larger for output growth, 12 times larger for the change in inflation, and 2 times larger for the change in the interest rate. Nonetheless, the structural matrices of the model are little changed relative to the baseline case. Consistent with this result, Figures 5–7 show that the main insights derived from the baseline specification are robust to allowing for heteroskedasticity in measurement errors.

## 4.4 Alternative Identification of the Temporary Monetary Shock

In the baseline formulation, the transitory monetary shock $z_t^m$ is identified by assuming that on impact it can change the policy rate but not output or inflation ($C_{12} = C_{22} = 0$). One potential criticism of this approach is that it may not be too reasonable at an annual frequency, as one year may be enough time for $z_t^m$ to affect output and inflation. To address this concern, we allow $z_t^m$ to have a nonpositive impact effect on these two variables. Specifically we impose the restrictions

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

If one is to allow for only one stationary monetary shock with these characteristics, it is necessary to impose the additional restriction that the impact effect on the policy rate of an increase in $z_t$—the other temporary shock in the model—be nonnegative, that is, one must impose

$$C_{34} \geq 0.$$

We implement this identification scheme by imposing appropriate restrictions on the prior distributions of the three parameters involved. Specifically, we assume that $-C_{12}$ and $-C_{22}$ have Gamma prior distributions with means 0.7 and 0.2, respectively, and standard deviations 0.2 and 0.1, respectively. We chose these values to match the impact effect on output and inflation of a monetary innovation after one year in the VAR model estimated in Christiano et al. (2005).[2] For the parameter $C_{34}$, we assume a uniform prior distribution over the interval $[0, 2]$.

As shown the bottom right panels of Figures 5 and 6 and the last row of Figure 7, the main results obtained from the baseline model—in particular, the short- and long-run effects of a fall in the natural rate on output, inflation, and interest rates—also hold under this alternative identification of the stationary monetary shock.

As noted in Uribe (2022), one need not restrict $C_{34}$ to be nonnegative to achieve an economically meaningful identification of both stationary shocks. Allowing $C_{34}$ to take negative and positive values opens the door for the existence of two monetary shocks. In this case, identification is achieved by the assumed differences in the prior distributions of the parameters of the second and fourth columns of $C$ and those of the second and fourth diagonal elements of the matrix $\rho$. For example, as explained in section 2.2, under the prior distribution $z_t^m$ is on average less persistent than $z_t$. We find (not shown) that the results of the paper also go through under this identification scheme when $C_{34}$ is assigned a standard normal prior distribution. As it turns out, the posterior means of $C_{34}$ and $C_{24}$ are positive, implying that $z_t$ is either a real shock or a monetary shock that does not depress output or inflation on impact in response to a policy tightening (e.g., a neo-Fisherian monetary shock).

# 5 Conclusion

What theories should we use to inform us about the economic consequences of persistent changes in consumers' subjective discount factors, the world real interest rate, or population growth? One way to make progress on these fundamental questions is to ask a related but simpler question: What happens with actual macroeconomic aggregates in the short and long runs in response to a change in the permanent component of the real interest rate? Having an answer to this stepping-stone question is important because it provides restrictions on

[2]Specifically, Figure 1 in that paper shows that a 0.75 percentage point fall in the policy rate causes 5 quarters later an increase in output of about 0.5 percent and an elevation in inflation of about 0.15 percentage points, with centered 95% error bands of approximate widths of 0.25 and 0.15 on each side (which we take to be approximately 2 standard deviations). Thus, we set the prior means of $C_{12}$ and $C_{22}$ at $0.7 \approx 0.5/0.75$ and $0.2 \approx 0.15/0.75$, respectively, and their prior standard deviations at $0.2 \approx 0.25/(0.75 \times 2)$ and 0.1 $\approx 0.15/(0.75 \times 2)$, respectively.

the class of theories of the natural rate of interest that are empirically sound.

In this paper, we formulate an empirical model with minimal identification restrictions suitable for estimating the macroeconomic effects of permanent disturbances to the real interest rate. We estimate the model on U.S. data over the period 1900 to 2021. The estimation results suggest that the answer to the question posed above falls on the side of a theory that predicts that a persistent fall in the natural rate of interest puts the economy on a lower trend trajectory and is contractionary and deflationary in the short run.

Importantly, the estimated model predicts that a permanent fall in the natural rate of interest depresses the trend of output not only when the economy is at the zero lower bound, but also when it is away from it. This result is intriguing because the most prominent if not the only theory suggesting that negative natural rate shocks can put the economy on a lower growth trajectory is the secular stagnation hypothesis, which relies on the economy being at the zero lower bound. A challenge for future research is therefore to develop a theoretical framework in which permanent falls in the natural rate lower potential output even when the policy rate is unconstrained.

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# WHAT DO LONG DATA TELL US ABOUT THE INFLATION HIKE POST COVID-19 PANDEMIC?

Stephanie Schmitt-Grohé
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**ABSTRACT**

To what extent is the recent spike in inflation driven by a change in its permanent component? We estimate a semi-structural model of output, inflation, and the nominal interest rate in the United States over the period 1900-2021. The model predicts that between 2019 and 2021 the permanent component of inflation rose by 51 basis points. If instead we estimate the model using postwar data (1955--2021), the permanent component of inflation is predicted to have increased by 238 basis points. A possible interpretation of this finding is that the model estimated on the shorter sample assigns a larger increase in the permanent component of inflation because the period 1955-2021 does not contain sudden sparks in inflation like the one observed in the aftermath of the COVID-19 pandemic but only gradual ones---the great inflation of the 70s took more than 10 years to build up. By contrast, the period 1900-1954 is plagued with sudden inflation hikes---including one around the 1918 Spanish flu pandemic---which the estimated model endogenously recalls and uses to interpret inflation around the COVID-19 episode. This result suggests that prewar data might be of use to understand recent inflation dynamics.

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# 1 Introduction

In the aftermath of the COVID-19 pandemic, the United States, like other advanced economies, experienced inflation at levels not seen for the past 40 years. In this note, we address a key question raised by this development. Namely, to what extent the recent spike in U.S. inflation is driven by a change in its permanent component.

To this end, we use a semi-structural model of output, inflation, and short-term interest rates to identify the permanent component of inflation. When we estimate the model over the period 1900 to 2021, it predicts a modest increase in the permanent component of inflation of 51 basis points between 2019 and 2021. By contrast, when we estimate the model using postwar data (1955 to 2021), the permanent component of inflation is predicted to have increased by 238 basis points between 2019 and 2021.

Our interpretation of this result is that the postwar data, dominated by the great inflation of the 1970s, which was slow in building up, leaves not much choice to the model but to interpret the 2021 inflation as also driven by low-frequency factors. By contrast, the pre-war era is rich in large and short-lived inflationary spikes, including the one around the Spanish flu pandemic of 1918. Since a key characteristic of the 2021 increase in inflation was its speed and size, the model estimated on long data naturally associates it more with the pre-war inflation spikes than with the great inflation of the 1970s.

We believe that this result is important for three reasons. First, most modern applied models of inflation dynamics are estimated on postwar data. The economic consequences of the COVID-19 crisis, however, are not alike anything we saw after World War II. As a consequence, it is conceivable that postwar data may contain little information useful for understanding economic outcomes of unusual events like the recent pandemic. Second, having a clear idea of the extent to which a given deviation of inflation from its intended target is driven by its permanent component is important for policymaking, as it informs the timing, size, and communication of the corresponding policy response. Third, it is becoming increasingly plausible that climate change will open an era of larger economic fluctuations. Until enough data has accumulated under such new regime, the volatile pre-war period may offer useful information for understanding and forecasting business cycles to come.

The remainder of this note proceeds in three sections. Section 2 presents the model and briefly discusses data and estimation. A more detailed presentation of these issues can be found in Schmitt-Grohé and Uribe (2022) from which this note draws heavily. Section 3 presents the main result of the paper, namely, the behavior of the permanent component of inflation around the COVID-19 pandemic and how it depends on whether the model is estimated on data starting at the beginning of the twentieth century or only after World

War II. Section 4 concludes.

## 2 The Empirical Model

The model structure is based on Schmitt-Grohé and Uribe (2022), which in turn builds on Uribe (2022). Here we briefly outline its key components. The log of real output per capita, $y_t$, the inflation rate, $\pi_t$, and the nominal interest rate, $i_t$ are assumed to be nonstationary. Output is assumed to be cointegrated with a nonstationary productivity shock, $X_t$, and with a nonstationary natural rate shock, $X_t^r$. Inflation is assumed to be cointegrated with a nonstationary inflation-target shock, $X_t^m$, and the nominal interest rate is assumed to be cointegrated with the nonstationary inflation-target shock and the natural rate shock. Accordingly, the cyclical components of output, inflation, and the nominal interest rate, denoted $\hat{y}_t$, $\hat{\pi}_t$ and $\hat{i}_t$, are given by

$$\hat{y}_t = y_t - X_t - \delta X_t^r,$$

$$\hat{\pi}_t = \pi_t - X_t^m,$$

and

$$\hat{i}_t = i_t - (1+\alpha)X_t^m - X_t^r,$$

where $\alpha$ and $\delta$ are estimated parameters.

The focus of the present paper is the behavior of the latent variable $X_t^m$ representing the permanent component of inflation.

The law of motion of the cyclical components is assumed to take the form

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

where $z_t^m$ and $z_t$ represent a stationary monetary shock and a stationary real shock, respectively. The matrices $B$ and $C$ are estimated.

The exogenous shocks follow univariate AR(1) processes,

$$\begin{bmatrix} \Delta X^m_{t+1} \\ z^m_{t+1} \\ \Delta X_{t+1} \\ z_{t+1} \\ \Delta X^r_{t+1} \end{bmatrix} = \rho \begin{bmatrix} \Delta X^m_{t} \\ z^m_{t} \\ \Delta X_{t} \\ z_{t} \\ \Delta X^r_{t} \end{bmatrix} + \Psi \begin{bmatrix} \epsilon^{X^m}_{t+1} \\ \epsilon^{z^m}_{t+1} \\ \epsilon^{X}_{t+1} \\ \epsilon^{z}_{t+1} \\ \epsilon^{X^r}_{t+1} \end{bmatrix}, \tag{1}$$

where $\rho$ and $\Psi$ are estimated diagonal matrices and $\epsilon^s_t$, for $s = X^m, z^m, X, z, X^r$, are i.i.d. disturbances distributed $N(0,1)$. All variables of the model are unobservable.

The model is estimated using data on output growth, $\Delta y_t$, the change in consumer price inflation, $\Delta \pi_t$, and the change in the short-term nominal interest rate, $\Delta i_t$. The consumer price index for a given year is computed as the arithmetic average of monthly observations. The following identities serve as observation equations:

$$\Delta y_t = \hat{y}_t - \hat{y}_{t-1} + \Delta X_t + \delta \Delta X^r_t + \mu^y_t,$$

$$\Delta \pi_t = \hat{\pi}_t - \hat{\pi}_{t-1} + \Delta X^m_t + \mu^\pi_t,$$

and

$$\Delta i_t = \hat{i}_t - \hat{i}_{t-1} + (1+\alpha)\Delta X^m_t + \Delta X^r_t + \mu^i_t,$$

where $\mu^s_t$, for $s = y, \pi, i$, are normally distributed mean-zero i.i.d. measurement errors whose variances are estimated.

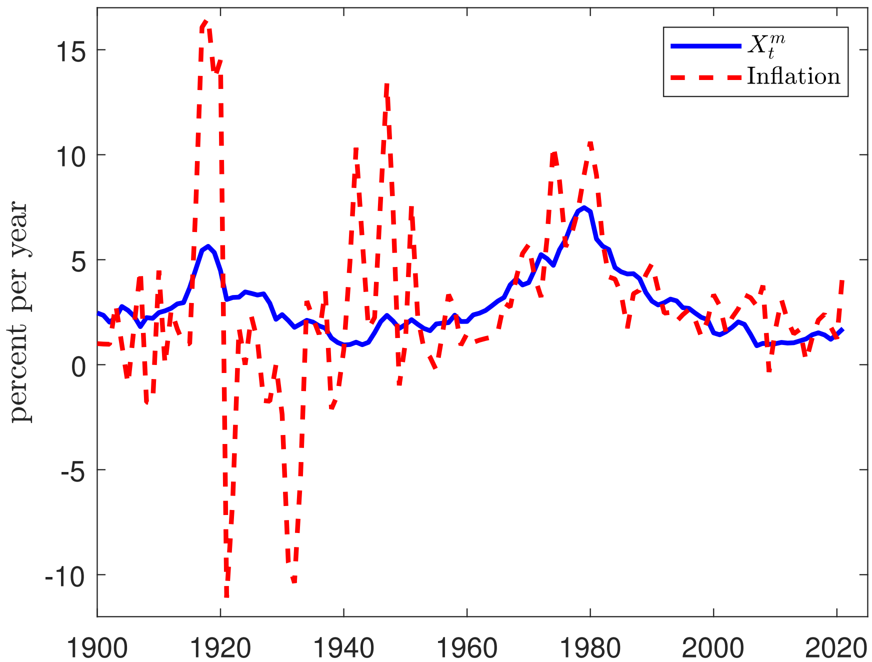
The model is estimated on annual U.S. data spanning the period 1900 to 2021. The data source for the period 1900 to 2017 is Jordá et al. (2017). Details on the data sources for the period 2018 to 2021, the identification scheme, priors, and the estimation technique can be found in Schmitt-Grohé and Uribe (2022).

# 3 The Permanent Component of Inflation

Figure 1 plots the estimated path of the permanent component of inflation, $X^m_t$, over the period 1900 to 2021, as predicted by the model. The path of $X^m_t$ was obtained by setting the vector of estimated parameters equal to its posterior mean and then smoothing using the Kalman filter. This technique yields $X^m_t$ up to a constant. In the figure, we arbitrarily pick this constant to ensure that the mean of $X^m_t$ over the estimation sample matches that of actual inflation. For comparison, the figure also plots the actual inflation rate.

Observed inflation displays distinct characteristics in the pre- and postwar periods. In the

Figure 1: Inflation and Its Permanent Component: Sample 1900 to 2021

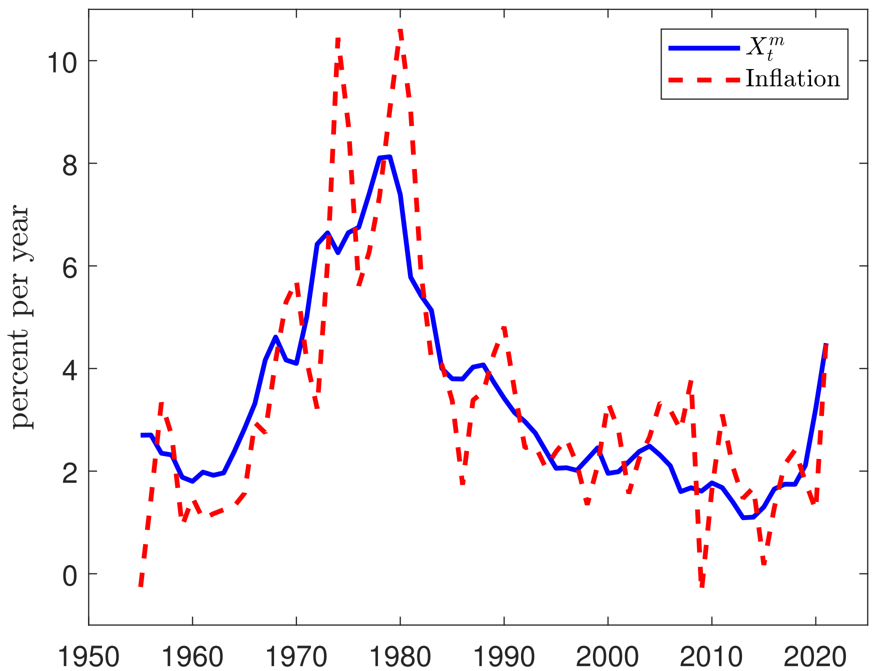



Notes. The permanent component of inflation, $X_t^m$, is computed by smoothing using the Kalman filter. It is normalized by adding a constant to match the sample mean of inflation.

pre-war period, inflation is highly volatile and spikes in inflation and deflation are typically short lived. As a result, the model interprets these episodes as mostly transitory, that is, not primarily driven by movements in the permanent component of inflation, $X_t^m$. A case in point is the observed inflation spike around the 1918 influenza pandemic. Between 1915 and 1918 inflation increased from 1 percent to 17 percent and then fell quickly to -11 percent by 1921. At the same time, $X_t^m$ was relatively little changed; it increased by 2 percentage points between 1915 and 1918 and then fell by 2.5 percentage points between 1918 and 1921.

Figure 1 further shows that the dynamics of inflation and its trend component are quite different in the postwar pre-COVID-19 era. This period is dominated by the great inflation of the 1970s. Contrary to what happened during the pre-war period, this inflation episode was slow in building up. Inflation began to increase from a level of 2 percent in the mid 1960s to a peak of 10 percent in 1980. The figure shows that inflation accelerated for about 15 years. The return of inflation to the levels observed in the mid 1960s took another six years. To a large extent, the model accounts for the great inflation of the 1970s with a significant movement in the permanent component of inflation, $X_t^m$. Between 1960 and 1980 the model estimates that $X_t^m$ rose 5 percentage points. In fact, the largest increase in $X_t^m$ over the entire 1900 to 2021 sample takes place during this episode.

Figure 2: Inflation and Its Permanent Component: Sample 1955 to 2021




Notes. The permanent component of inflation, $X_t^m$, is computed by smoothing using the Kalman filter. It is normalized by adding a constant to match the sample mean of inflation.

A key characteristic of the post-COVID-19 inflation hike is its speed. From 2019, the year before the onset of the pandemic, to 2021, the annual rate of inflation rose by 277 basis points. The model interprets this sudden spike in inflation as being more akin to those observed in the pre-war period than to the great inflation of the 1970s. Specifically, of the 277 basis point increase in inflation observed between 2019 and 2021, $X_t^m$ accounts for only 51 basis points. Thus, according to the model the 2021 inflation burst was not predominantly driven by an increase in the permanent component of inflation.

How important is the inclusion of pre-war data in arriving at this conclusion? This question is relevant because most of the existing literature on the joint behavior of inflation, output, and the nominal interest uses data that starts only after World War II. We therefore next examine the predictions of the model when estimated on postwar data. Specifically, we estimate the model on a sample starting in 1955. In keeping with much of the related literature, we start the postwar sample a few years after the actual end of World War II. Figure 2 displays the inferred path of the permanent component of inflation, $X_t^m$, and actual inflation. As in the estimate using data since 1900, the inflation of the 1970s is interpreted by the model as having a large permanent component. The key difference of this estimate relative to the one obtained when the model is estimated over the sample beginning in

1900 is that now almost all of the COVID-19 inflation spike is attributed to the permanent component of inflation $X_t^m$. Specifically, between 2019 to 2021 the permanent component of inflation increases by 238 basis points, which is more than eighty percent of the total observed inflation increase of 277 basis points. So according to the model estimated on postwar data, the inflation burst associated with COVID-19 was to a large extend driven by the permanent component of inflation.

This result is robust to allowing for heteroskedasticity in measurement errors. Specifically, if in the estimation using data from 1900 to 2021 measurement errors are allowed to explain a fraction of the variance of the data twice as large pre 1955 than post 1955, then of the 277 basis-point actual increase in inflation between 2019 and 2021, the model attributes 89 basis points to the permanent component $X_t^m$. While this is larger than the 51 basis points obtained in the absence of heteroskedasticity, it is still only about 1/3 of the 238 basis points attributed to $X_t^m$ when the model is estimated on the 1955–2021 sample.

Which of the two interpretations of the inflation hike of 2021 makes more economic sense? Neither the model nor the data can answer this question. We believe, however, that the more compelling view is the one that emerges from the estimation including the pre-war data. The reason is that the economic developments triggered by the COVID-19 pandemic are of a nature not seen since the end of World War II. So it is conceivable that postwar data does not have that much to say about such event. By contrast, the pre-1955 period was littered with economic crises that came with large swings in the rate of inflation, including a pandemic similar to the COVID-19 one. It seems therefore reasonable that data from that early period may provide useful information for understanding the economic predicament in which the economy found itself in the aftermath of the global health crisis caused by the COVID-19 pandemic.

Including the pre-war period in the estimation of the model also matters for the forecasted level of the permanent component of inflation. Figure 3 displays $X_t^m$ for the period 2015 to 2021 (solid lines) and its forecast for the out-of-sample period 2022 to 2030 (dotted lines) under the two estimations considered, namely, 1900 to 2021 and 1955 to 2021. In the figure the level of $X_t^m$ is normalized to 2 percent in 2019. The figure shows that, as discussed earlier, when the model is estimated on the shorter sample, $X_t^m$ is predicted to have increased during COVID-19 more than four times as much as when the model is estimated on the longer sample. This difference in the importance attributed to the permanent component of inflation is also present in out of sample forecasts. The model assumes that the change in the permanent component of inflation follows a univariate first-order autoregressive process (see equation 1). For the estimate based on data from 1900 to 2021, and using the mean of the posterior estimate of $\rho$ and the smoothed value for $\Delta X_{2021}^m$, the out-of-sample forecast

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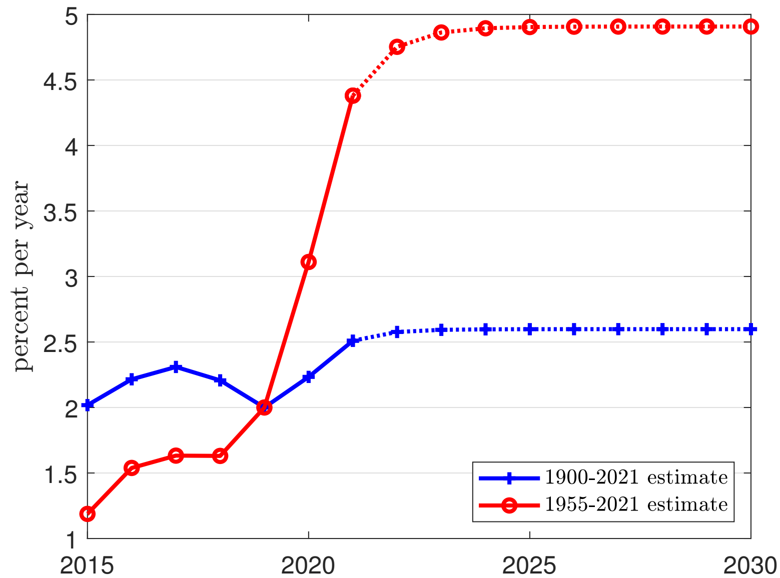


Notes. Forecasts are computed using the mean of the posterior parameter distribution. For 2015 to 2021, the smoothed values of $X_t^m$, normalized to 2 percent in 2019, are shown and for 2022 to 2030 the forecasted values.

for $X_t^m$ by 2030 is only 9 basis points higher than its 2021 value. For the estimate of the model based on the shorter sample, $X_t^m$ is forecasted to be 53 basis points above its 2021 value by 2030. This difference is explained mainly by the difference in the estimated values of $\Delta X_{2021}^m$ rather than in the estimated values of $\rho$.

In sum, the estimates based on the long sample imply that by 2030 the permanent component of inflation will exceed its 2019 value by 60 basis points whereas estimates based on the short sample suggest that by 2030 the permanent component of inflation will exceed its 2019 level by 291 basis points. This means that according to the estimates based on the long sample the inflation spike that arrived with the COVID-19 pandemic is largely temporary whereas according to the estimates based on the short sample it is largely permanent.

# 4 Conclusion

Modern business cycle analysis is a story told with postwar data. Before the COVID-19 pandemic this approach made sense. The volatile pre-war data seemed out of touch with the unprecedented stability witnessed in the postwar period and in particular since the Great Moderation. This note suggests that the COVID-19 pandemic has called this approach into question and has given renewed value to the information contained in pre-war macroeconomic indicators. It is now the postwar data that seem out of touch with current developments.

The analysis conducted in this note serves as a proof of concept. Seen from the perspective of a model estimated on postwar data, the 2021 inflation spur is interpreted to be caused by a large increase in the permanent component of inflation. The reason is that the model was given little chance to conclude otherwise; the only other major prior inflation increase during this sample period turned out to be a protracted one, which naturally is ascribed to the permanent component—not just by the present model but also by the majority of existing models of U.S. inflation. However, once the sample is expanded to include the sudden, large, and short-lived swings in inflation observed in the first half of the 20th century, the same model attributes a major fraction of the post COVID-19 inflation to its transitory component.

Looking ahead, recent studies point in the direction that climate change will raise economic volatility around the globe. Thus, until sufficient data accumulates under this seemingly emerging new regime, long historical time series data could be a valuable input to business cycle analysis.

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