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THE NEO-FISHER EFFECT IN THE UNITED STATES AND JAPAN

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ABSTRACT

I investigate the effects of an increase in the nominal interest rate on inflation and output in the United States and Japan during the postwar period. I postulate a structural autoregressive model that allows for transitory and permanent nominal and real shocks. I find that nominal interest-rate increases that are expected to be temporary, lead, in accordance with conventional wisdom, to a temporary increase in real rates that is contractionary and deflationary. By contrast, nominal interest-rate increases that are perceived to be permanent cause a temporary decline in real rates with inflation adjusting faster than the nominal interest rate to a higher permanent level. Estimated impulse responses show that inflation reaches its long-run level within a year. Importantly, because real rates are low during the transition, the economy does not suffer an output loss. This result is relevant for the design of monetary policy in economies plagued by chronic below-target inflation, for it is consistent with the prediction that a credible announcement of a gradual return of nominal rates to normal levels can bring about a swift convergence of inflation to its target level without negative consequences for aggregate activity.

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1 Introduction

What is the effect of an increase in the nominal interest rate? One can argue on theoretical grounds that the answer to this question depends on whether the increase in the interest rate is expected to be permanent or transitory. Viewed through the lens of the new-Keynesian model with a central bank that observes the Taylor principle, 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 current inflation. By contrast, if the increase in the nominal interest rate is perceived to be permanent, sooner or later, inflation will have to increase by roughly the same magnitude, since the real interest rate, given by the difference between the nominal rate and expected inflation, is unlikely to be determined by nominal factors in the long run. This one-to-one long-run relationship between nominal rates and inflation is known as the Fisher effect and is built into most modern dynamic macroeconomic models, including those in the new-Keynesian tradition. 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? Recent theoretical work argues that a credible permanent increase in the nominal interest rate causes an immediate increase in inflationary expectations. This result has come to be known as the neo-Fisher effect. In this paper, I investigate the short-run effects of permanent nominal-interest-rate shocks from an empirical perspective.

I postulate a structural vector autoregressive (SVAR) model in three endogenous variables, output, inflation, and the nominal interest rate. The model is driven by four disturbances: a permanent monetary shock, a transitory monetary shock, a permanent nonmonetary shock, and a transitory nonmonetary shock. To identify these four driving forces, I impose a number of restrictions on the structure of the model: First, I assume that both inflation and the nominal interest rate are cointegrated with the permanent monetary shock and share a common cointegrating vector. This assumption implies that the Fisher effect

holds in the long run. Section 2 provides some evidence in support of this assumption. Second, I assume that output is cointegrated with the permanent nonmonetary shock. Finally, I assume that temporary increases in the nominal interest rate have a nonpositive impact effect on output and inflation. I estimate the SVAR model on postwar quarterly data from the United States and Japan using Bayesian techniques.

The estimated SVAR model predicts that a transitory increase in the nominal interest rate produces dynamics that are in line with the conventional wisdom: the real interest rate increases on impact and converges from above to its steady-state, with depressed levels of aggregate activity and inflation. By contrast, the estimated SVAR model predicts that in response to a permanent increase in the nominal interest rate, the one-to-one adjustment in the inflation rate predicted by the Fisher effect happens in the short run, by which I mean less than a year. Furthermore, inflation rises faster than the nominal interest rate. As a result, the real interest rate falls on impact and converges from below to its steady state. In line with this effect on real rates, the SVAR model predicts that the adjustment to a permanent increase in the nominal interest rate does not generate a loss of aggregate output.

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 theoretical front, Schmitt-Grohé and Uribe (2014 and 2017) show that the neo-Fisher effect obtains in the context of standard dynamic optimizing models with flexible or rigid prices, respectively. Specifically, they show that a credible increase in the nominal interest rate that is expected to be sustained for a prolonged period of time can give rise to an immediate increase in inflationary expectations. Cochrane (2017) shows that a standard macroeconomic model coupled with an equilibrium selection criterion that avoids interest-rate jumps delivers neofisherian dynamics in which the nominal interest rate and inflation positively comove in the short run. Cochrane (2015) presents an nontechnical exposition of the neo Fisher effect and uses it to shed light on the role of monetary policy during the Great Contraction of 2007

and its aftermath. Erceg and Levin (2003) study a calibrated dynamic general equilibrium model with nominal rigidity in which private agents have imperfect information about the permanent and transitory components of monetary-policy shocks. They show that imperfect information of this type can provide an adequate explanation of the observed inflation persistence during disinflation episodes. To my knowledge, there are no econometric studies of the neo-Fisher effect. However, there is a related empirical literature devoted to estimating long-run movements in the parameters describing monetary policy, including the inflation target, and their economic effects, to which I now refer in a non-exhaustive manner. Sims and Zha (2006) estimate a regime-switching model for U.S. monetary policy and find that during the postwar period there were three policy regime switches, but that they were too small to explain the observed increase in inflation of the 1970s or the later disinflation that started with the Volker chairmanship. Ireland (2007) estimates a new-Keynesian model with a time-varying inflation target and shows that, possibly as a consequence of the Fed's attempt to accommodate supply-side shocks, the target increased significantly during the 1960s and 1970s and fell sharply in the early 2000s. Using a similar framework, Milani (2009) shows that movements in the inflation target become less pronounced if one assumes that agents must learn about the level of the inflation target. De Michelis and Iacoviello (2016) estimate an SVAR model with permanent inflation-target shocks to evaluate the Japanese experience with Abenomics. Finally, Fève, Matheron, and Sahuc (2010) study a dynamic optimizing model with persistent inflation-target shocks and 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 remainder of the paper is presented in 6 sections. Section 2 presents evidence consistent with the long-run validity of the Fisher effect. Section 3 presents the SVAR model. Section 4 discusses the observables used in the estimation, the assumed prior distributions of the estimated parameters, and the estimation procedure. The main results of the paper

are contained in sections 5 and 6, which present the estimated effects of permanent and temporary interest-rate shocks in the United States and Japan, respectively. Section **??** closes the paper with a discussion of actual monetary policy in the ongoing low-inflation era from the perspective of the estimated model.

2 The Fisher Effect

My analysis of the neo-Fisher effect assumes the empirical validity of the Fisher effect, interpreted as a description of the long-run relationship between the nominal interest rate and inflation. In this section, I present some empirical evidence consistent with this relationship. The Fisher equation takes the form

$$i_t = R_t + E_t \pi_{t+1}$$

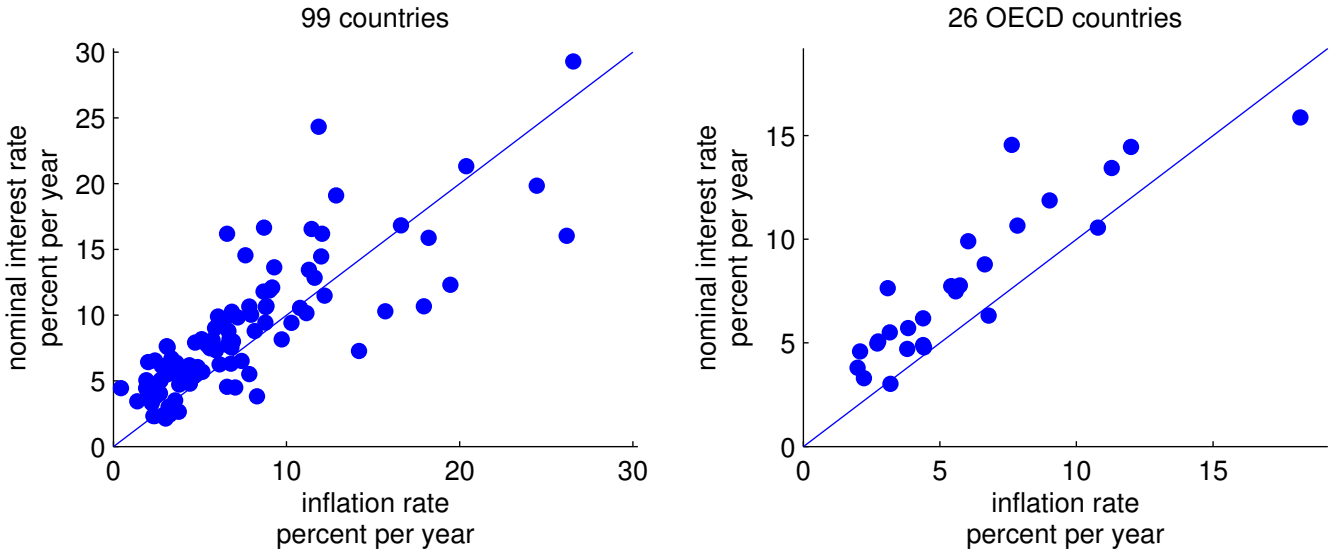
where i_t denotes the nominal interest rate, R_t denotes the real interest rate, π_t denotes the inflation rate, and E_t denotes expectations conditional on information available in period t. This expression says that the nominal interest rate incorporates two types of compensation to lenders, one is a compensation for the loss of purchasing power of money due to expected inflation during the investment period, and the other is a real retribution for the sacrifice of present consumption. Assuming that in the long run expected inflation equals actual inflation, we have that

$$i = R + \pi,$$

where variables without a subscript refer to long-run values. Further assuming that in the long run the real interest rate is determined solely by real factors (such as technology, distortionary taxes, or economic openness), the above expression delivers a one-to-one long-run relationship between the nominal interest rate and the rate of inflation.

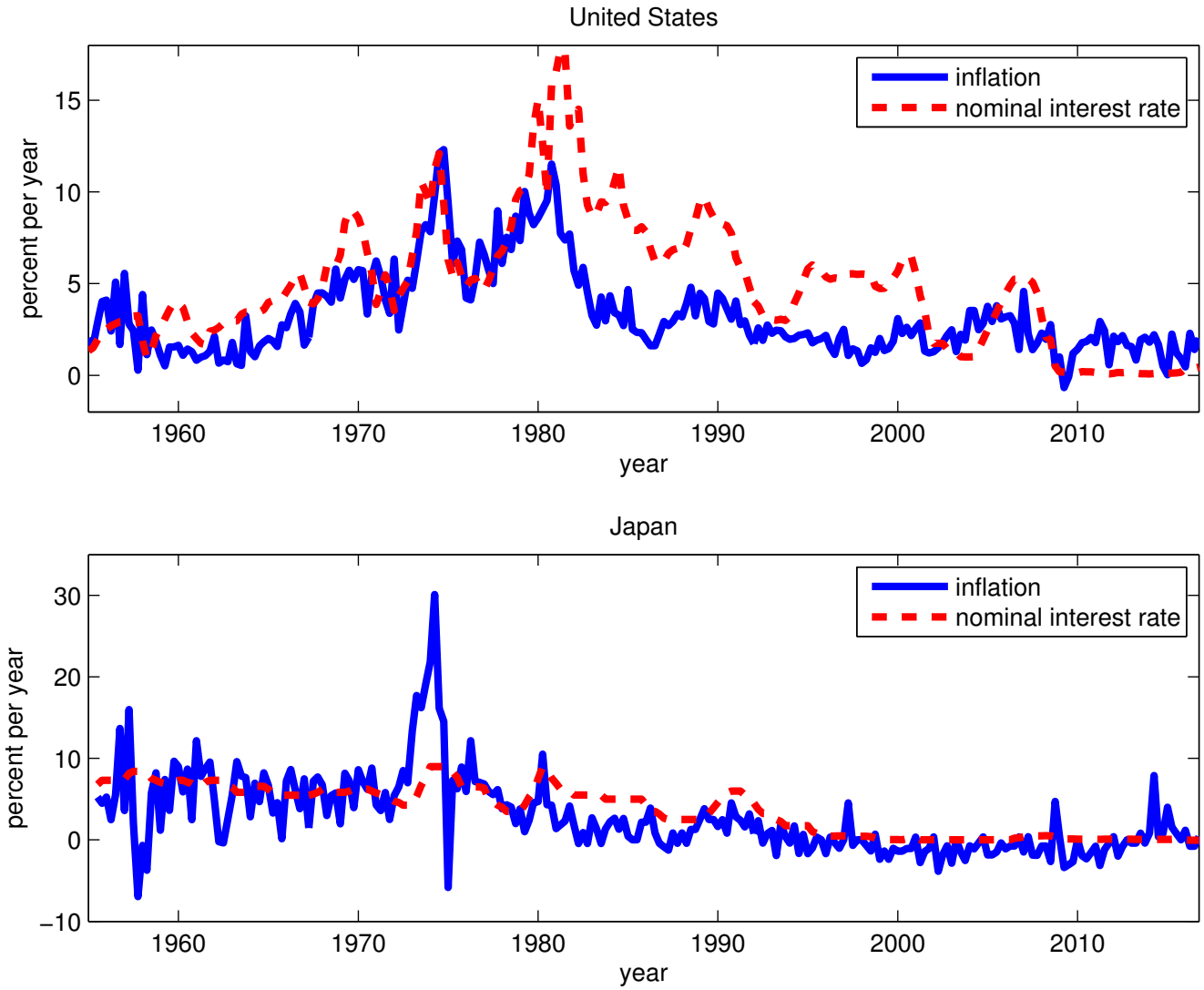
Figure 1 displays average rates of inflation and nominal interest rates across countries.

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



Notes. Each dot represents one country. For each country, averages are taken over the longest available noninterrupted sample. The average sample covers the period 1989 to 2012. The solid line is the 45-degree line. Source: World Development Indicators (data.worldbank.org/indicator). Inflation is the CPI inflation rate (code FP.CPI.TOTL.ZG). The nominal interest rate is the t-bill rate, computed as the difference between the lending interest rate (code FR.INR.LEND) and the risk premium on lending (lending rate minus treasury bill rate, code FR.INR.RISK). Countries for which one or more of these series were missing as well as outliers, defined as countries with average inflation or interest rate above 50 percent, were dropped from the sample.

Figure 2: Inflation and the Nominal Interest Rate in the United States and Japan



Notes. Quarterly frequency. Source: See section 5 for the United States and section 6 for Japan.

The average sample covers the period 1989 to 2012. The scatter plots look consistent with the Fisher effect, in the sense that on average increases in the interest rate are roughly associated with one -for-one increases in the rate of inflation. This is the case for all the countries in the sample as well as for the subsample of OECD countries, which are on average half as inflationary as the group of no-member 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 and Japan

over the period 1955 to 2016. 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. In both countries, the high-inflations of the 1970s and 1980s coincided with high levels of the interest rate. Symmetrically, the disinflations that took place in both countries since the 1990s were accompanied by low nominal rates.

Having provided some evidence consistent with the assumption that the Fisher effect holds in the long run, I now turn to the central focus of my investigation, the identification of permanent and temporary nominal-interest-rate shocks and their dynamic effects on inflation, output, and the real interest rate.

3 The SVAR Model

The empirical model aims to capture the dynamics of three macroeconomic indicators, namely, the logarithm of real output per capita, denoted y_t, the inflation rate, denoted π_t and expressed in percent per year, and the nominal interest rate, denoted i_t and also expressed in percent per year. Let Y_t be a vector collecting these three variables,

$$Y_t \equiv \begin{bmatrix} y_t \\ \pi_t \\ i_t \end{bmatrix}.$$

I assume that Y_t is 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^n, and a stationary nonmonetary shock, denoted z_t^n. As mentioned earlier, the focus of my analysis is to compare the short-run effects of permanent and transitory interest-rate shocks, embodied in the exogenous variables X_t^m and z_t^m. The shocks X_t^n and z_t^n are meant to capture the nonstationary and stationary components of combinations of nonmonetary disturbances of different natures, such as tech-

nology shocks, preference shocks, or markup shocks, which my analysis is not intended to individually identify.

Let $\tilde{Y}_t$ be a vector containing detrended output, detrended inflation, and the detrended nominal interest rate. Formally,

$$\tilde{Y}_t \equiv \begin{bmatrix} (y_t - X_t^n) \times 100 \\ \pi_t - X_t^m \\ i_t - X_t^m \end{bmatrix}.$$

Because the shocks X_t^m and X_t^n are not observable, neither is $\tilde{Y}_t$. Let $\hat{Y}_t$ denote the deviation of $\tilde{Y}_t$ from its unconditional mean, that is,

$$\hat{Y}_t \equiv \begin{bmatrix} \hat{y}_t \\ \hat{\pi}_t \\ \hat{i}_t \end{bmatrix} = \tilde{Y}_t - E\tilde{Y}_t,$$

where E denotes the unconditional expectations operator. I assume that $\tilde{Y}_t$ is stationary. This means that output is cointegrated with X_t^n and that inflation and the nominal interest rate are both cointegrated with X_t^m. Because inflation and the nominal interest rate share a common nonstationary component, they are cointegrated. In other words, I am assuming that the Fisher effect holds, in the sense that shocks that cause a permanent change in the nominal interest rate also cause the same permanent change in the inflation rate. But the assumption that $\hat{Y}_t$ is stationary says nothing about the neo-Fisher effect, that is, about the short-run effect on inflation and output of a permanent change in the interest rate.

I assume that the law of motion of $\hat{Y}_t$ takes the autoregressive form

$$\hat{Y}_t = \sum_{i=1}^{L} B_i \hat{Y}_{t-i} + C u_t \tag{1}$$

where

$$u_t \equiv \begin{bmatrix} x_t^m \\ z_t^m \\ x_t^n \\ z_t^n \end{bmatrix},$$

$$x_t^m \equiv \Delta X_t^m - \Delta X^m$$

and

$$x_t^n \equiv (\Delta X_t^n - \Delta X^n) \times 100.$$

with Δ denoting the time-difference operator, $\Delta X^m \equiv E\Delta X_t^m$, and $\Delta X^n \equiv E\Delta X_t^n$. The variables x_t^m and x_t^n denote demeaned changes in the nonstationary shocks. The objects B_i, for $i = 1, \ldots, L$, are 3-by-3 matrices of coefficients, C is a 3-by-4 matrix of coefficients, and L is a scalar denoting the lag length of the SVAR system.

I assume that u_t follows an AR(1) law of motion of the form

$$u_{t+1} = \rho u_t + \psi \epsilon_{t+1}, \tag{2}$$

where ρ and ψ are 4-by-4 diagonal matrices of coefficients, and ϵ_t is a 4-by-1 i.i.d. disturbance distributed $N(\emptyset, I)$.

Thus far, I have introduced three identification assumptions, namely, that output is cointegrated with X_t^n and that inflation and the interest rate are cointegrated with X_t^m. In addition, I impose the following restrictions

$$C_{12} \leq 0, \quad C_{22} \leq 0, \text{ and } C_{31} \geq -1,$$

where C_{ij} denotes the (i, j) element of C. The first two of these conditions restrict transitory exogenous increases in the interest rate to have nonpositive impact effects on output or inflation. The restriction on C_{31} implies that a permanent exogenous increase in the nominal

interest rate has a nonnegative impact effect on the nominal interest rate itself. Finally, I introduce the normalizations $C_{32} = C_{14} = 1$.

4 Observables, Priors, and Estimation Method

To estimate the unobservable system (1)-(2), I use observations of output growth expressed in percent per quarter, the change in the nominal interest rate, and the interest-rate-inflation differential, defined as

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

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

$$\begin{aligned} 100 \times \Delta y_t &= 100 \times \Delta X^n + \hat{y}_t - \hat{y}_{t-1} + x_t^n \\ r_t &= r + \hat{i}_t - \hat{\pi}_t \\ \Delta i_t &= \Delta X^m + \hat{i}_t - \hat{i}_{t-1} + x_t^m \end{aligned} \tag{3}$$

where $r \equiv Er_t$ represents the unconditional mean of the interest-rate-inflation differential. I assume that Δy_t, r_t, and Δi_t are observed with measurement error. Formally, letting o_t be the vector of variables observed in quarter t, I assume that

$$o_t = \begin{bmatrix} \Delta y_t \times 100 \\ r_t \\ \Delta i_t \end{bmatrix} + \mu_t \tag{4}$$

where μ_t is a 3-by-1 vector of measurement errors distributed i.i.d. $N(\emptyset, R)$, where R is a diagonal variance-covariance matrix.

I estimate the model on quarterly data using Bayesian techniques. To compute the

likelihood function, it is convenient to use the state-space representation of the model. Let

$$\xi_t \equiv \begin{bmatrix} \hat{Y}_t \\ \hat{Y}_{t-1} \\ \vdots \\ \hat{Y}_{t-L+1} \\ u_t \end{bmatrix}.$$

Then the system composed of equations (1), (2), (3), and (4) can be written as follows:

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

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

where the matrices F, P, A, and H are known functions of B_i, $i = 1, \ldots L$, C, ρ, ψ, ΔX^n, ΔX^m, and r and are presented in the appendix. This representation allows for the use of the Kalman filter to evaluate the likelihood function (see, for example, Hamilton, 1994, chapter 13).

I consider two lag specifications, 4 and 8 quarters, which are the lag lengths used in the vast majority of the related literature. Prominent examples of each specification are Christiano, Eichenbaum, and Evans (2005) who estimate an SVAR model with stationary monetary and nonmonetary shocks with 4 lags, and Blanchard and Quah (1989), who estimate the effects of stationary and nonstationary shocks in an SVAR model with 8 lags.

Table 1 displays the prior distributions of the estimated coefficients. The prior distributions of all elements of B_i, for $i = 1, \ldots, L$, are assumed to be normal. In the spirit of the Minnesota prior (MP), I assume a prior parameterization in which at the mean the elements of $\tilde{Y}_t$ follow univariate autoregressive processes. So when evaluated at the prior mean, only the main diagonal of B_1 takes nonzero values and all other elements of B_i for $i = 1, \ldots, L$ are nil. Because the system (1)-(2) is cast in terms of stationary variables, I deviate from

Table 1: Prior Distributions

Parameter	Distribution	Mean.	Std. Dev.
Main diagonal elements of B_1	Normal	0.95	0.5
All other elements of B_i, $i = 1, \ldots, L$	Normal	0	0.25
C_{21}	Normal	-1	1
$-C_{12}, -C_{22}, 1 + C_{31}$	Gamma	1	1
All other estimated elements of C	Normal	0	1
ψ_{ii}, $i = 1, 2, 3, 4$	Gamma	1	1
ρ_{ii}, $i = 1, 2, 3$	Beta	0.3	0.2
ρ_{44}	Beta	0.7	0.2
$100 \times \Delta X^n, r, \Delta X^m$	Normal	$\text{mean}(o_t)$	$\sqrt{\frac{\text{var}(o_t)}{T}}$
R_{ii}	Uniform $\left[0, \frac{\text{var}(o_t)}{10}\right]$	$\frac{\text{var}(o_t)}{20}$	$\sqrt{\frac{\text{var}(o_t)}{120}}$

Note. T denotes the sample length.

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. As in the MP, I impose higher prior standard deviations on the diagonal elements of B_1 (0.5) than on the remaining elements of B_i (0.25) for $i = 1, \ldots, L$.

The coefficient C_{21} takes a normal prior distribution with mean -1 and standard deviation 1. The value assigned to the mean of this distribution 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. All other unrestricted parameters of the matrix C are assigned a normal prior distribution with mean 0 and standard deviation 1.[1] The remaining 3

[1]One might wonder whether a rationale like the one I used to set the prior mean of C_{21} could apply to C_{13}, the parameter governing the impact output effect of a nonstationary nonmonetary shock, X_t^n, 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 $y_t = X_t^n + z_t^n + \alpha k_t + (1 - \alpha)h_t$ 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^n. In this case, the contemporaneous effect of a unit increase in X_t^n 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^n 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 not clear, giving credence to a prior of 0 for C_{13}. One could further think about the role of variable input utilization. An increase in X_t^n is likely to cause an increase in utilization, favoring a prior mean of 0 over one of -1 for C_{13}.

estimated elements of C are, as explained above, subject to inequality restrictions. Specifically, $-C_{12}$, $-C_{22}$, and $1+C_{31}$ are restricted to be nonnegative. I assume that these objects have Gamma prior distributions with mean and standard deviations equal to one.

The parameters ψ_{ii}, for $i = 1, \ldots, 4$, representing the standard deviations of the four exogenous innovations in the AR(1) process (2) are all assigned Gamma prior distributions with mean and standard deviation equal to one. I impose nonnegative serial correlations on the four exogenous shocks ($\rho_{ii} \in (0, 1)$ for $i = 1, \ldots, 4$), and adopt beta prior distributions for these parameters. I assume relatively small means of 0.3 for the prior serial correlations of the two monetary shocks and the nonmonetary nonstationary shock and a relatively high mean of 0.7 for the stationary nonmonetary shock. The small prior mean serial correlations for the monetary shocks reflect the usual assumption in the related literature of serially uncorrelated monetary shocks. The relatively small prior mean serial correlation for the nonstationary nonmonetary shock reflects the fact that the growth rate of the stochastic trend of output is typically estimated to have a small serial correlation. Similarly, the relatively high prior mean of the serial correlation of the stationary nonmonetary shock reflects the fact that typically these shocks (e.g., the stationary component of TFP) are estimated to be persistent. The prior distributions of all serial correlations are assumed to have a standard deviation of 0.2.

The unconditional means of the three observables are assumed to have normal prior distributions with means equal to their sample means and standard deviations equal to their sample standard deviations divided by the square root of the length of the sample period. Finally, the variances of all measurement errors are asumed to have a uniform prior distribution with lower bound 0 and 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 an Monte-Carlo Markov chain of one million draws after burning the initial 100 thousand 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 thousand with replacement.

5 The Neo-Fisher Effect in the United States

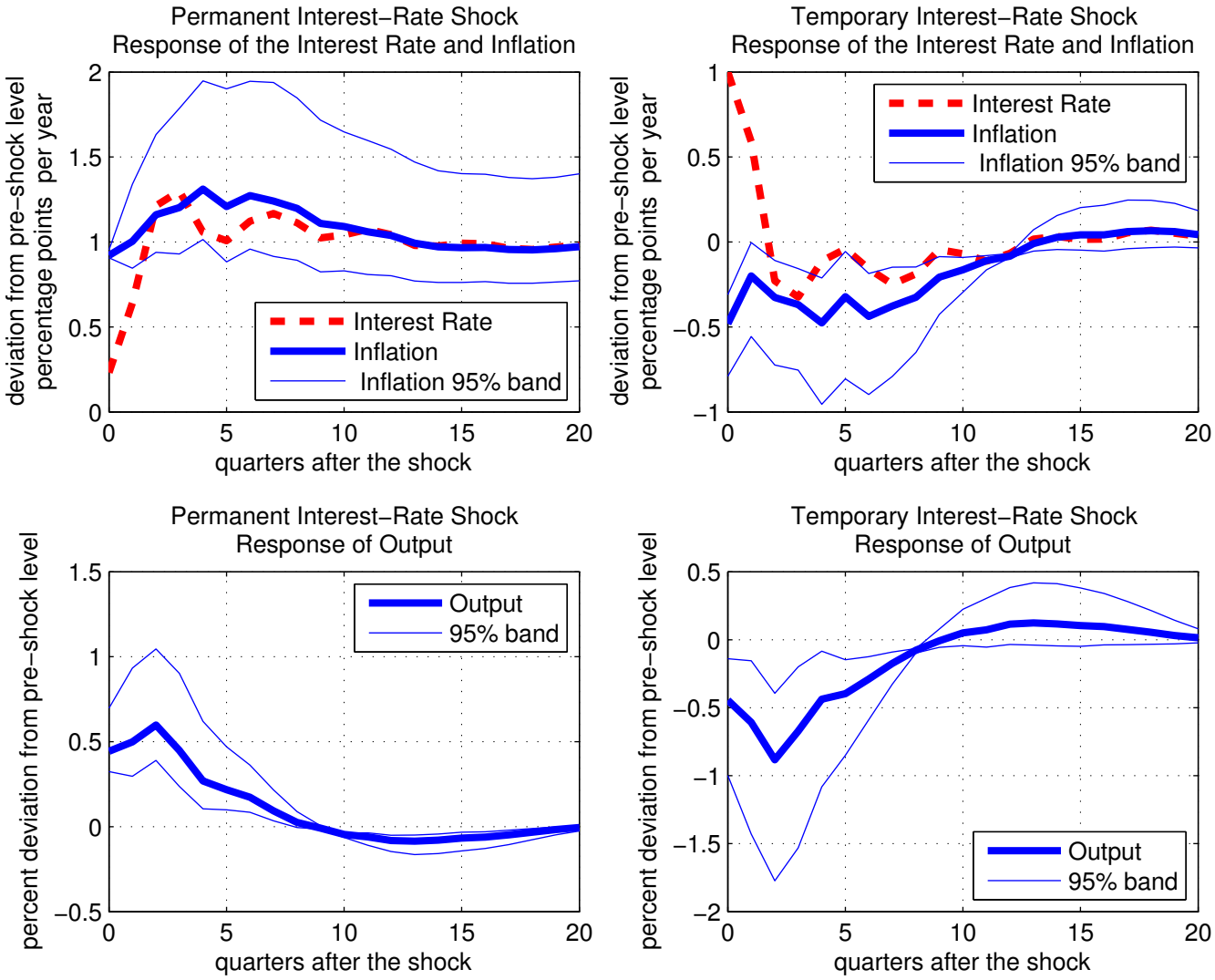
For the United States, I use quarterly data spanning the period 1954:Q3 to 2016:Q4. The proxy for y_t is the logarithm of real GDP seasonally adjusted in chained dollars of 2009 minus the logarithm of the civilian noninstitutional population 16 years old or older. The proxy for π_t is the growth rate of the implicit GDP deflator expressed in percent per year. In turn, the implicit GDP deflator is constructed as the ratio of GDP in current dollars and real GDP both seasonally adjusted. The proxy for i_t is the monthly Federal Funds Effective rate converted to quarterly frequency by averaging and expressed in percent per year. The source for nominal and real GDP is the Bureau of Economic Analysis (bea.gov), the source for population is the Bureau of Labor Statistics (bls.gov) and the source for the Federal Funds rate is the Board of Governors of the Federal Reserve System (federalreserve.gov).

The identification scheme assumes that the interest rate, inflation, and output, all possess unit roots, and that the interest rate and inflation are cointegrated. Before plunging in the predictions of the SVAR model, it might be of some interest to briefly report the empirical relevance of these assumption from the perspective of univariate representations of the data. Specifically, the augmented Dickey-Fuller (ADF) test, which is a commonly used test of the null hypothesis of a unit root, fails to reject the null hypothesis for y_t, i_t, and π_t, and rejects it for $i_t - \pi_t$ at standard confidence levels of 10 percent or less.[2]

Figure 3 displays mean posterior estimates of the responses of inflation, output, and the nominal interest rate to permanent and temporary interest-rate shocks, along with asymmetric 95-percent error bands constructed using the method proposed by Sims and Zha (1999).

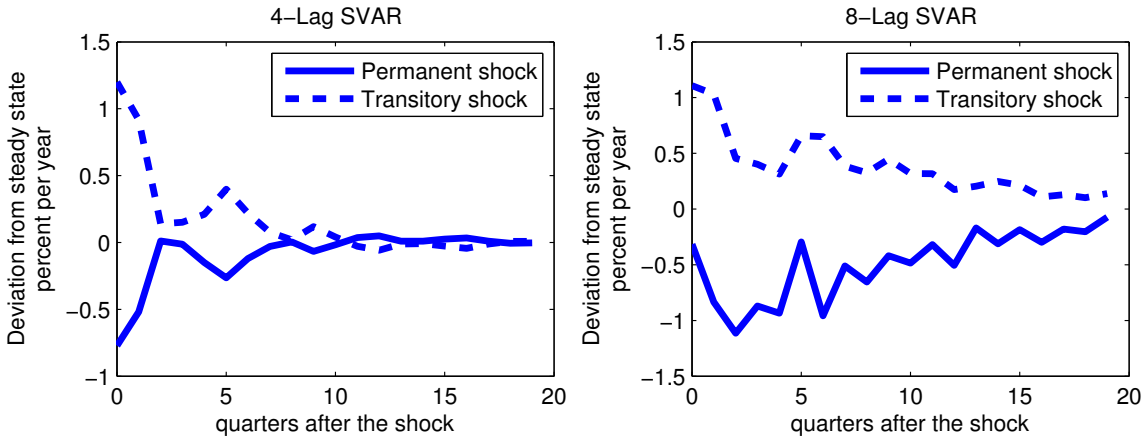
[2]Specifically, for a random variable x_t, the ADF test considers the null hypothesis that $x_t = x_{t-1} + \eta_0 + \sum_{i=1}^{I} \eta_i \Delta x_{t-i} + \epsilon_t$, where ϵ_t is white noise, against the alternative hypothesis that $x_t = \delta x_{t-1} + \gamma t + \eta_0 + \sum_{i=1}^{I} \eta_i \Delta x_{t-i} + \epsilon_t$, with $\delta < 1$. For i_t and π_t, I restrict γ to be zero (no time trend), and for $i_t - \pi_t$ I restrict $tgamma$ and η_0 to be zero (no time trend or drift). I include 4 lags of Δx_t ($I = 4$). The p values for $x_t = y_t, i_t, \pi_t, i_t - \pi_t$ are, respectively, 0.705, 0.154, 0.142, and 0.0365.

Figure 3: Impulse Responses to Interest-Rate Shocks: United States



Notes. The SVAR model includes 4 lags. Impulse responses are posterior mean estimates. Asymmetric error bands are computed using the Sims-Zha (1999) method.

Figure 4: Response of the Real Interest Rate to Permanent and Transitory Interest-Rate Shocks



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

The size of the permanent interest-rate shock is set to ensure that on average it leads to a 1 percent increase in the nominal interest rate in the long run, where the average is taken over the posterior distribution of impulse responses. 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 the adjustment of inflation to its higher long-run level takes place in the short run. In fact, inflation increases by 1 percent on impact and remains around that level thereafter.

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. The left panel of 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 0.8 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.

By contrast, the responses of nominal and real variables to a transitory interest-rate

shock, shown in the two right-side 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 the left panel of 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 interest-rate shock.

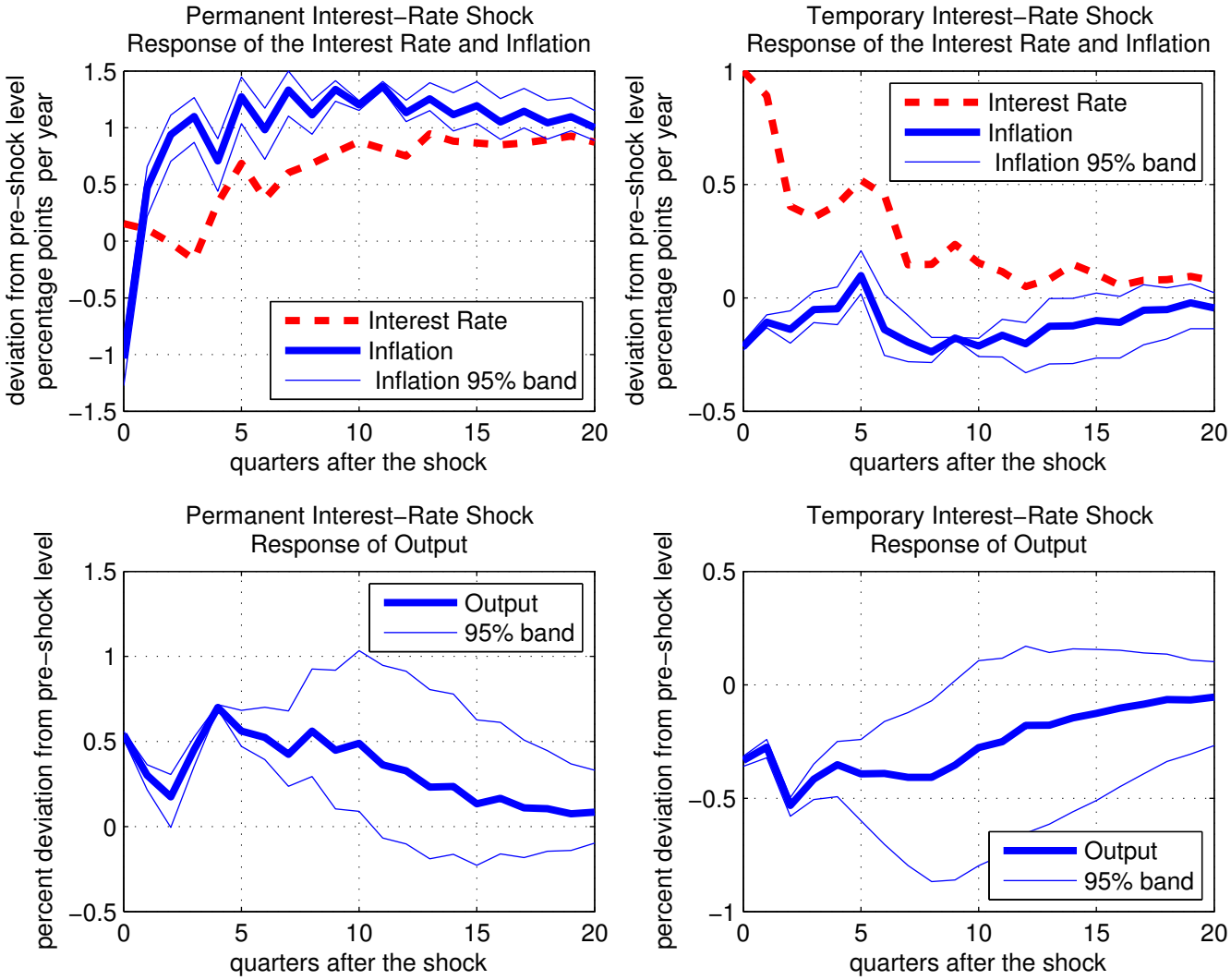
Figure 5 displays impulse responses predicted by an SVAR specification that includes 8 lags of the endogenous variables. All other aspects of the model and its estimation are as before. The general message of the figure is the same as under the 4-lag specification, that is, following a permanent increase in the nominal interest rate, the rate of inflation rises to its long-run level in the short run. A difference with respect to the 4-lag specification is that now reflation does not occur instantaneously, but takes one quarter to materialize. In the quarter in which the interest-rate shock takes place (period 0), inflation falls by one annualized percentage point. However, just one quarter later (period 1) inflation is 0.5 annualized percentage points above its pre-shock level and in the second quarter after the shock (period 2) it reaches 0.93 percentage points above its pre-shock level, which is virtually equal to its long-run value. As in the 4-lag SVAR specification, the response to a permanent interest-rate shock is characterized by an expansion in aggregate activity and low real interest rates (figure 4).

6 The Neo-Fisher Effect in Japan

I now use Japanese data to estimate the SVAR model developed in section 3. The prior distributions continue to be the ones displayed in table 1. The data is quarterly and spans the period 1955:Q2 to 2016:Q4. Real GDP and the GDP deflator are taken from Japan's Cabinet Office (esri.cao.go.jp).[3] Population is taken from Official Statistics of Japan (e-stat.go.jp) and

[3]The source presents the data in three subsamples, 1955:Q2-1999:Q1, 1980:Q1-2011:Q2, and 1994:Q1-2017:Q1, which requires concatenation. The GDP deflator for the first subsample appears not to be seasonally adjusted (no information in English is provided by the source in this regard). For this reason, this series was seasonally adjusted using the software x-13arima-seats, developed by the United States Census Bureau. In

Figure 5: Impulse Responses to Interest-Rate Shocks: United States, Eight-Lag SVAR Specification



Note.See notes to figure 3.

comes at an annual frequency. A quarterly series was obtained by geometric interpolation. The proxy for the nominal interest rate is the discount rate until the second quarter of 1995 and the call rate from the third quarter of 1995 until the end of the sample, reflecting the Japanese monetary authority's choice of the policy instrument. The source for both series is the Bank of Japan (stat-search.boj.or.jp) and both are monthly. A quarterly series was obtained by averaging. Given the series for real GDP, the GDP deflator, population, and the nominal interest rate, the proxies for the observables used in the estimation of the SVAR model, Δy_t, r_t, and Δi_t, were constructed as in the case of the United States.

ADF tests fail to reject the null hypothesis that y_t and i_t have a unit -root and reject the hypothesis that $i_t - \pi_t$ has a unit root at standard significance levels, which is consistent with the assumed identification scheme, but does reject the hypothesis that π_t has a unit root.[4]

Although the sample begins in 1955, I use 1975:Q1 as the starting point of the baseline estimation, because the growth rate of real GDP per capita displays a significant break around this time. Figure 6 displays the level and growth rate of real GDP per capita over the period 1955 to 2016. It is clear from the figure that Japan experiences a significant slowdown in the trend of GDP per capita around 1975, possibly marking the end of the postwar reconstruction. The average growth rate of GDP per capita from 1955 to 1975 was more than three times as large as over the period 1975 to 2016 (7.03 versus 1.98 percent per year). The first subperiod was also more volatile, with a standard deviation of real per capita GDP growth of 6.04 percent compared to 4.08 percent over the more recent subperiod. This marked change in the statistical properties of output growth can affect the estimated cyclical properties of the SVAR model, justifying the exclusion of the earlier subsample. Notwithstanding this problem, at the end of this section, I conduct an estimation of the SVAR model over the entire sample, as a robustness check.

the third subsample, the observation 2017:Q1 was dropped, as it was preliminary when it was downloaded.

[4]The p values for y_t, i_t, π_t, and $i_t - \pi_t$ are 0.414, 0.549, 0.032, and 0.001, respectively. These tests are robust to starting the sample in 1975:Q1, with associated p values of 0.843, 0.306, 0.027, and 0.021. See footnote 2 for more details.

Figure 6: Real GDP Per Capita: Japan 1955:Q3 to 2016:Q4

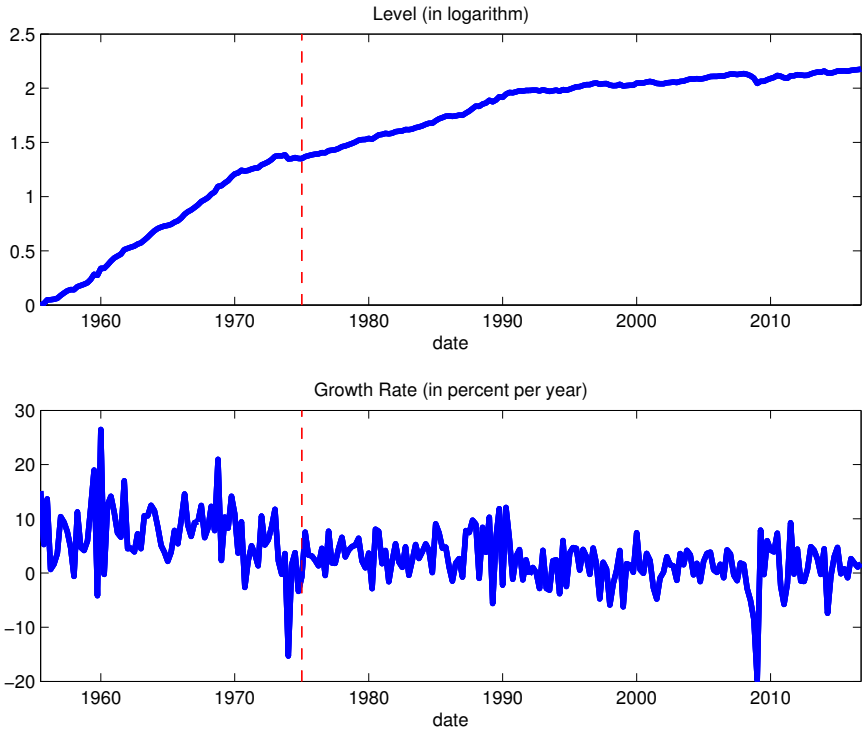
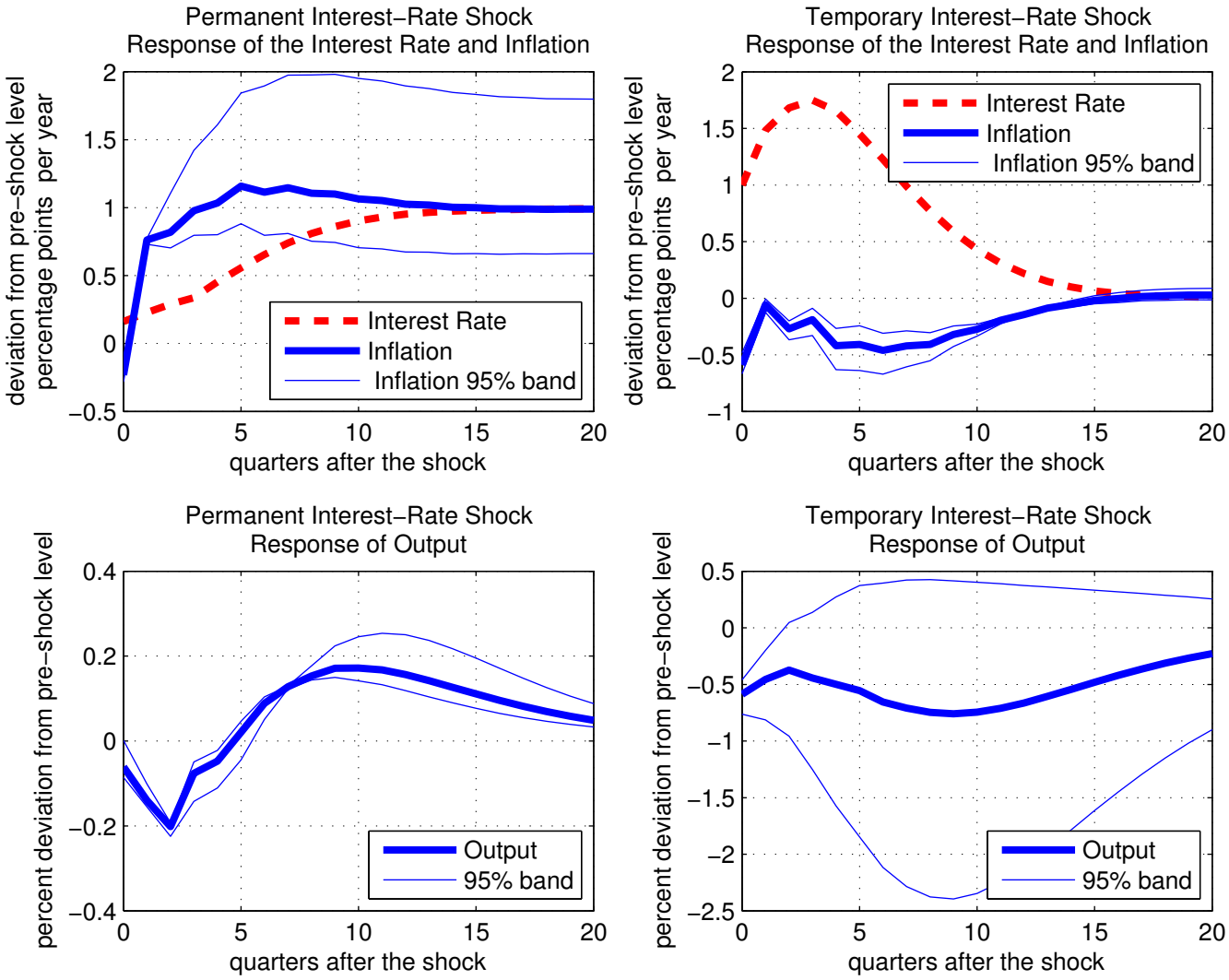


Figure 7 displays impulse responses of inflation, the nominal interest rate, and output to permanent and transitory interest-rate shocks. The construction of the figure is as explained in section 5. As in the case of the estimation on U.S. data. In response to a permanent increase in the interest rate of 1 percent, the economy reflates quickly. One quarter after the shock, the inflation rate is 0.75 percent above its pre-shock level, and three quarters after the shock it is 0.98 percent above its pre-shock level, virtually at its long-run position. Output experiences a short contraction, but recovers quickly after period 2, rising and staying above trend afterward, with a net output gain throughout the adjustment process. By contrast, a transitory increase in the interest rate causes falls in output and inflation, with both variables converging from below to their respective steady states.

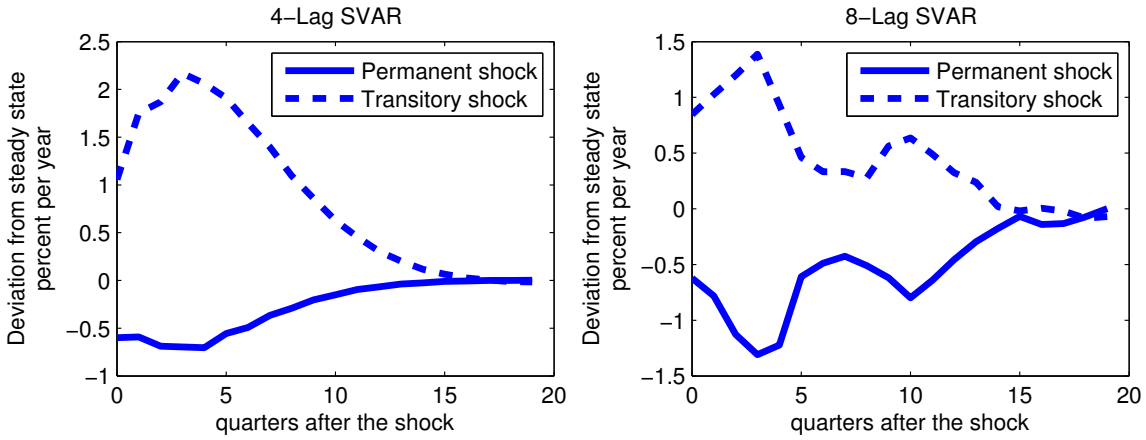
The response of the real interest rate to permanent and transitory interest-rate shocks follows the same pattern identified in the estimates using U.S. data. In particular, as figure 8 shows, a permanent increase in the interest rate is associated with a fall in the real interest rate and a gradual convergence from below, exactly the opposite of what happens after a

Figure 7: Impulse Responses to Interest-Rate Shocks: Estimates on Japanese Data



Notes. The SVAR model includes 4 lags. See notes to figure 3.

Figure 8: Response of the Real Interest Rate to Permanent and Transitory Interest-Rate Shocks: Estimates on Japanese Data

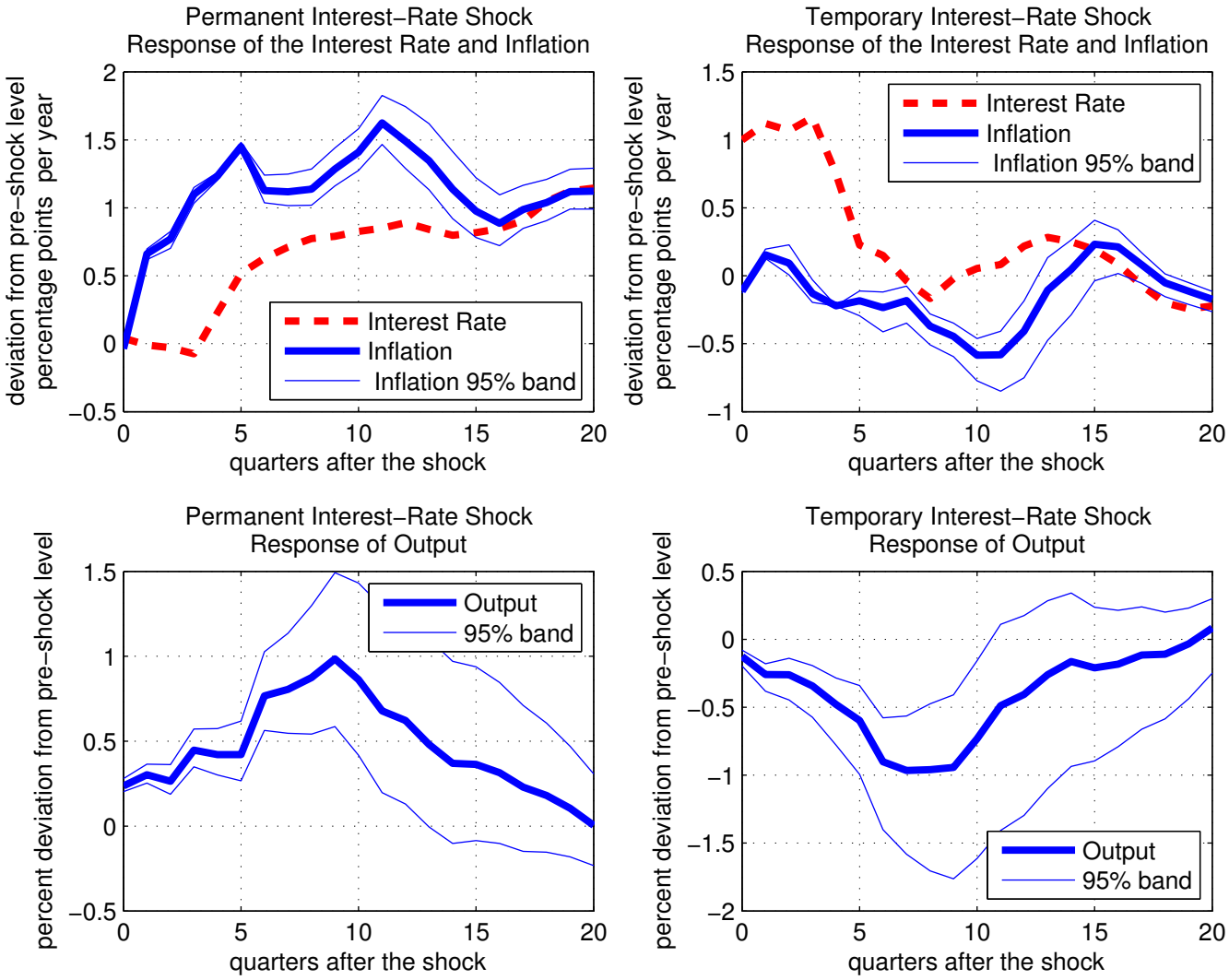


Note. See note to figure 4.

temporary increase in the interest rate, which triggers a rise in the real interest rate and convergence from above.

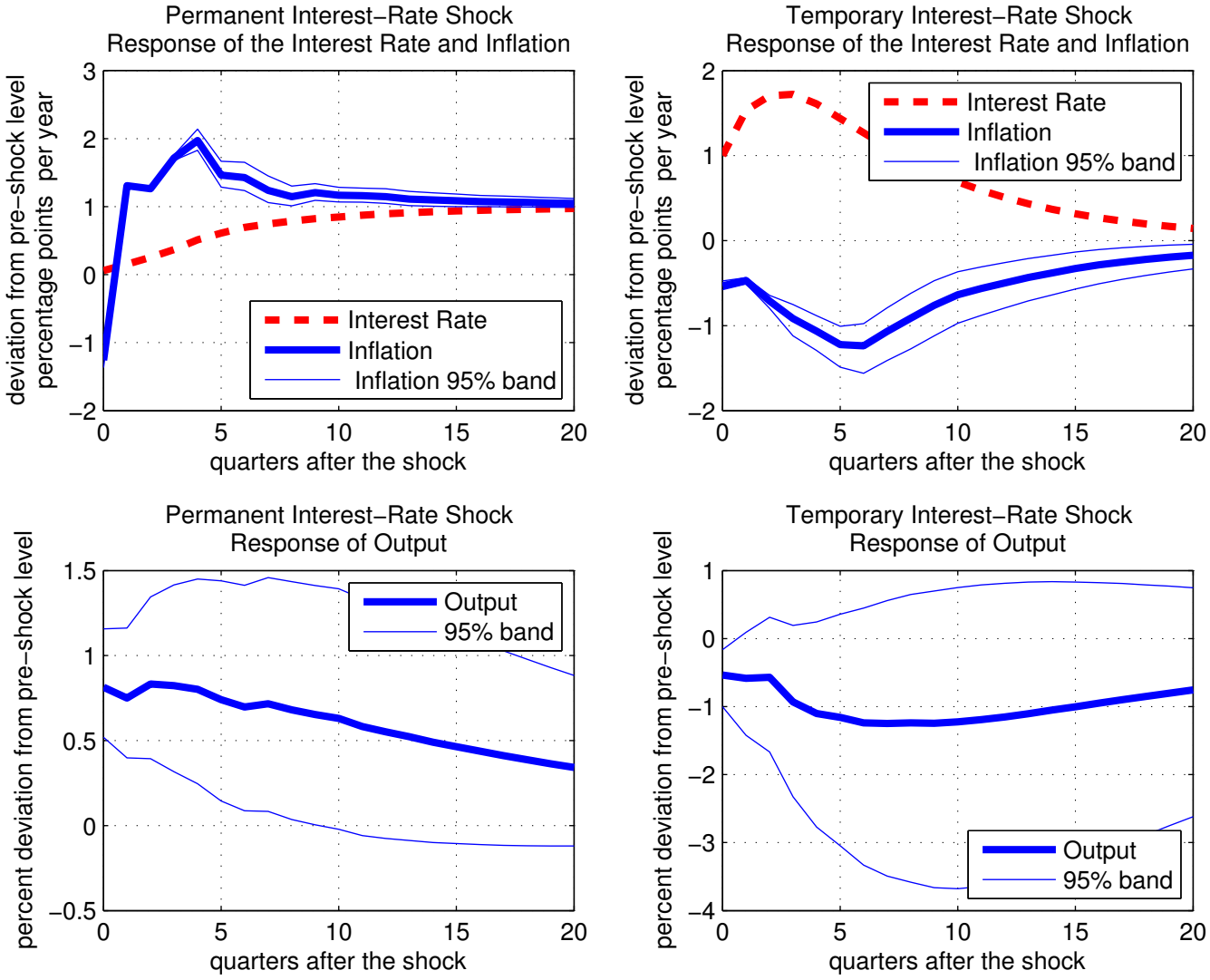
I close this section with two robustness checks. The first one, illustrated in figure 9 extends the lag length of the SVAR model from 4 to 8 quarters. The main result of the paper obtains, namely, a permanent increase in the interest rate is accompanied by a swift convergence of the inflation rate to its higher permanent level. Importantly, the small and short-lived contraction in output implied by the 4-lag version of the model is not present in the 8-lag specification, suggesting that a recession caused by a permanent increase in the nominal interest rate is not a robust prediction of the SVAR model estimated on Japanese data. Finally, figure 10 displays the predictions of the SVAR model estimated on Japanese data over the period 1955:Q3 to 2016:Q4. The quick reflation of the economy following the announcement of a long-run increase in interest rates also obtains over the long sample. As expected, given the analysis of the break in the output trend discussed earlier in this section, the model estimated on the long sample has difficulties identifying the output effect of monetary shocks, which is reflected in wide error bands. Nonetheless, the point estimates indicate that a permanent increase in the nominal interest rate is not contractionary.

Figure 9: Impulse Responses to Interest-Rate Shocks: Eight-Lag SVAR Model Estimated on Japanese Data



Note.See notes to figure 3.

Figure 10: Impulse Responses to Interest-Rate Shocks: Estimates on Japanese Data from 1955:Q3 to 2016:Q4



Note.See notes to figure 3.

7 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 this type of monetary shock transmits 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. Soon enough these economies find themselves with zero or negative nominal rates and with the low-inflation problem not going away. After some time, the Fisher effect kicks in, and 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 Fisher effect).

In this paper I argue, based on econometric evidence drawn from U.S. and Japanese data, that a gradual and permanent increase in the nominal interest rate causes a fast adjustment of inflation to a permanently higher level, low real interest rates, and no output loss. These findings are consistent with the neo-Fisherian prediction that a credible announcement of a gradual return of the nominal interest rate to normal levels can achieve a swift reflation of the economy with sustained levels of economic activity.

Appendix

Matrices of the State-Space Representation

Let

$$B \equiv [B_1 \cdots B_L],$$

and let I_j denote an identity matrix of order j, $\emptyset_j$ denote a square matrix of order j with all entries equal to zero, and $\emptyset_{i,j}$ denote a matrix of order i by j with all entries equal to zero. Also let L, S, and V denote, respectively, the number of lags, the number of shocks, and the number of endogenous variables included in the SVAR model. Then, for $L \geq 2$ we have

$$F = \begin{bmatrix} B & C\rho \\ \left[I_{V(L-1)}\, \emptyset_{V(L-1),V}\right] & \emptyset_{V(L-1),S} \\ \emptyset_{S,VL} & \rho \end{bmatrix}, P = \begin{bmatrix} C\psi \\ \emptyset_{V(L-1),S} \\ \psi \end{bmatrix};$$

$$A' = \begin{bmatrix} 100 \times \Delta X^n \\ r \\ \Delta X^m \end{bmatrix}, \text{ and } H' = \begin{bmatrix} M_\xi & \emptyset_{V,V(L-2)} & M_u \end{bmatrix},$$

where, in the specification considered in the body of the paper ($S = 4$, $V = 3$, and a particular ordering of the endogenous and exogenous variables in the vectors $\hat{Y}_t$ and u_t), the matrices M_ξ and M_u take the form

$$M_\xi = \begin{bmatrix} 1 & 0 & 0 & -1 & 0 & 0 \\ 0 & -1 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & -1 \end{bmatrix} \text{ and } M_u = \begin{bmatrix} 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 \end{bmatrix}.$$

The case $L = 1$ is a special case of $L = 2$ in which $B_2 = \emptyset_V$.

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