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Inflation Dynamics When Inflation is Near Zero

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Abstract:
This paper discusses the likely evolution of U.S. inflation in the near and medium term on the basis of (1) past U.S. experience with very low levels of inflation, (2) the most recent Japanese experience with deflation, and (3) recent U.S. micro evidence on downward nominal wage rigidity. Our findings question the view that stable long-run inflation expectations and downward nominal wage rigidity will provide sufficient support to prices such that deflation can be avoided. We show that an inflation model fitted on Japanese data over the past 20 years, which accounts for both short- and long-run inflation expectations, matches the recent U.S. inflation experience quite well. While the model indicates that U.S. inflation might be subject to a lower bound, it does not rule out a prolonged period of mild deflation going forward. In addition, our micro evidence on wages suggests no obvious downward rigidity in the firm's wage costs, downward rigidity in individual wages notwithstanding. As a consequence, downward nominal wage rigidity may provide little offset to deflationary pressures in the current U.S. situation, despite some circumstantial evidence that this channel might have been at work in the past.

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1. Introduction

In the wake of the longest postwar recession in U.S. history, inflation declined noticeably from its 2008 peak (see table 1 for changes in a variety of inflation measures). At about a 1 percent annual rate at the time of writing (the start of 2011), U.S. inflation is as low as it has been since the early 1960s. With most economists estimating that significant economic slack remains in the American economy, it is reasonable to wonder about the future trajectory of inflation. Will well-anchored expectations pull inflation up as the recovery proceeds? Will slack resources pull inflation toward or below zero? Is the United States headed for a Japanese-style period of protracted, albeit modest deflation? Will downward nominal wage rigidity provide an offset to disinflationary forces?

A key difficulty in answering these questions is that the United States has had very little recent experience with very low inflation rates. A number of researchers have examined the behavior of inflation over the past 20 years, as it fell from an annual rate noticeably above 2 percent to a yearly average of about 2 percent from the late 1990s through the mid-2000s. But this paper is concerned with how inflation behaves as it drops *below* 2 percent, and the U.S. experience in this range is quite limited. One might wish to look for a breakpoint in the time-series properties of inflation data in recent years, but it is nearly impossible to detect a breakpoint in the last few years of a series.[1]

While there are many possible approaches to exploring these questions, we pursue a few specific tacks. We consider a relatively wide array of Phillips curve specifications. In several instances these specifications trade off micro-foundations and/or rational expectations for empirical relevance, and throughout the paper we highlight the similarities and differences with the more standard models. First we examine the period from 1954 to 1963, a time when U.S. inflation was low. The evidence from this early episode suggests a relatively benign outcome for current inflation developments, a result of well-anchored long-run inflation expectations and a very mild tradeoff between inflation and resource slack when slack is sizable. This latter effect could be the result of downward nominal wage rigidities. Yet the rest of the paper questions whether stable long-run inflation expectations can prevent deflation and whether downward nominal wage rigidity can provide an important offset to price deflation in the current environment. For this purpose, we

[1] While not reported here, we also conduct conventional unknown breakpoint tests for changes in the autoregressive properties of a variety of U.S. inflation measures. Simple autoregressive and reduced-form models of the core and headline inflation measures examined in table 1 above typically find either one or no breaks in the series from 1970 to the present, with the breakpoints falling in the period between 1975 and 1980.

analyze the parallels between Japan's long experience with low inflation over the last two decades and the relatively recent experience in the United States. We then examine some preliminary disaggregated data on wages to explore the extent to which firms are able to adjust their wage costs downward; if there is downward rigidity in the firm's wage bill, this may act to limit the fall in inflation that might otherwise occur.

The relatively short samples we are examining make any inflation forecasting exercise even more difficult than usual. Nevertheless, our findings cast doubt on the now fairly widespread notion that stable long-run inflation expectations will provide sufficient support to prices to avoid a prolonged period of low inflation, or even deflation. Indeed, recent U.S. inflation dynamics do not seem to differ much from the Japanese experience of the post-1990s. While in Japan a deflationary spiral has not materialized, there is also little in the data that would lead us to confidently rule out the possibility of a prolonged period of mild deflation in the United States. In addition, our microeconomic evidence on wages suggests that downward nominal wage rigidity may provide little offset to deflationary price pressures. This lack of a counterweight results from the fact that downward nominal wage rigidity in individual wages may not necessarily translate into significant downward nominal rigidity in firms' wage bills.

The rest of the paper is structured as follows. Section 2 examines an optimistic scenario for near-term inflation developments in the United States. In this scenario, inflation is anchored by long-run inflation expectations and the correlation between real activity and inflation is nonlinear, so that a large degree of resource slack has a disproportionately smaller influence on inflation than does a more modest amount of slack. This "optimistic" model, based on estimating a stable nonlinear relationship between inflation and the output gap in the 1950s and early 1960s, predicts U.S. inflation bottoming out at a value slightly below 1 percent and, therefore, no deflation. The rest of the paper examines more closely the two key features of this simple model, namely the role of long-run inflation expectations as a driver of inflation and downward nominal wage rigidity as a potential buffer to declining inflation. Specifically, section 3 examines the Japanese experience, with an emphasis on assessing the role that inflation expectations at different horizons (short- and long-run) play in driving inflation. Section 4 draws parallels between Japanese and U.S. inflation dynamics over the past two decades. On the basis of this comparative analysis, section 5 reaches more pessimistic conclusions about the outlook for U.S. inflation in the near- and medium-term. Section 6 examines micro wage data, focusing on preliminary evidence about the presence of downward rigidity (or lack thereof) in the firm's overall wage bill. Section 7 provides some concluding remarks.

We consider different specifications of inflation because U.S. episodes with very low levels of inflation in the post-WWII period are scant. Therefore, statistical tests for choosing one model over another have limited power. Moreover, sources of instability in the inflation process have been documented extensively in the literature. These instabilities have made inflation difficult to forecast over time. It is also possible that the dynamics of inflation at 2 percent or lower will turn out to differ fundamentally from past experiences with low inflation. For all of these reasons, it is important to acknowledge model uncertainty under current circumstances, and broaden the analysis to more than just a single inflation model.

2. An Optimistic Scenario: A Nonlinear Trend-Inflation Model of U.S. Inflation

In this section we revisit the one U.S. postwar episode (other than the most recent period) when inflation was very low. From 1954 to 1963,[2] annual consumer price inflation averaged about 1.6 percent and fluctuated within a relatively narrow range. The private consumption expenditures deflator's maximum year-over-year increase over this period was 3.25 percent. The conduct of monetary policy in these years was reevaluated in favorable terms by Romer and Romer (2002). After examining the FOMC records of the time, Romer and Romer claim that the Federal Reserve "showed the same overarching concern about inflation that is the hallmark of post-Paul Volcker monetary-policy orthodoxy." In particular, they argue that policy tightening in response to increases in expected inflation was more aggressive than in the late 1960s and in the 1970s. Moreover, the 1954-1963 episode is potentially interesting not just because inflation was low, but also because there were some relatively large fluctuations in real activity.

It is apparent from figure 1 that over the period 1954:Q1 to 1963:Q4 there were some run-ups in inflation during expansions, followed by declines in inflation immediately after recessions. This pattern is consistent with a Phillips curve-type relationship. Figure 2 plots an estimate from the Congressional Budget Office (CBO) of the output gap against the four-quarter-ahead inflation rate, as measured by the PCE deflator.[3] While not especially tight, the data show a positive relationship between inflation and the output gap. The data also suggest the presence of a nonlinearity, with inflation increasing at a faster rate for a given increase in the output gap when the output gap becomes more positive. This nonlinearity mimics, *mutatis mutandis*, the one found by Phillips (1958)

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[2] We exclude the early 1950s from the analysis because of the 1951 and 1952 price control amendments associated with the Korean War.

[3] The CBO output gap in this section's figures is depicted as a four-quarter moving average with declining weights to smooth out some short-term volatility.

in his original study of the relationship between wage inflation and the unemployment rate in the United Kingdom. In the U.S. context, such a feature of the data is the result of a few quarterly observations and, as a result, it should be taken with caution.

The observations over the period from 1954 to 1963 suggest a modest tradeoff between inflation and resources slack overall, and especially so when there is a large amount of slack. This is confirmed by estimating the following relationship between inflation and the output gap:

$$\ln(1+0.01*(\pi_{t+4}^{4}-\pi_{t}^{LR})) = \underset{(0.0118)}{0.0464} - \underset{(0.0071)}{0.0278}*\ln(5.3-0.50*\tilde{y}_t), \quad n=40,\ R^2=0.46, \tag{2.1}$$

where standard errors for the estimates are in parentheses.[4] In the equation, π_{t+4}^{4} is the four-quarter-ahead inflation rate, while $\tilde{y}$ denotes the output gap.[5] The inflation rate is expressed as a deviation from long-run inflation expectations, π^{LR}, as of time *t*. Since survey expectations of inflation for this period are not available, we assume them to be constant and equal to the average annual rate of inflation (1.6 percent) prevailing over the period. We apply a linear transformation to the output gap before taking logs, as the output gap can take negative values.[6] This linear transformation can be thought of as providing an unemployment rate equivalent of the output gap by means of a simple Okun's law.[7] The estimated constant in equation (2.1) is constrained to yield inflation equal to its long-run expected rate when the output gap is zero.

The estimated log-relationship (2.1) allows the tradeoff between inflation and resource slack to vary at different levels of the output gap. This specification is very similar to the relationship considered in Phillips's original article. The estimated nonlinearity in U.S. data over the period 1954 to 1963 is fairly small. An increase in the output gap from 2 to 4 percent raises annual inflation by seven-tenths of 1 percent (0.70). The same improvement in the output gap, but from –4 to –2

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[4] Standard errors are corrected to account for a moving average component in the estimated error term.

[5] In equation (2.1), four-quarter-ahead inflation is computed as: $\pi_{t+4}^{4} = 100*\left(\frac{JC_{t+4}}{JC_t}-1\right)$, where JC is the total PCE deflator. The output gap is: $\tilde{y}_t = 100*\ln(Y_t / Y_t^*)$, where Y_t is the actual GDP level and Y_t^* is the CBO estimate of potential GDP.

[6] For the same reason, a linear transformation before taking logs is also applied to the difference between inflation and long-run expected inflation.

[7] The constant in the linear transformation (which we set to 5.3) can be interpreted as the equilibrium unemployment rate, while the value of the slope (set at 0.5) implies that a positive output gap of 1 percent yields an unemployment rate 50 basis points below the natural rate. These values are consistent with the estimates obtained from a simple level Okun's law relationship over the sample 1954:Q1 to 1963:Q4.

percent, raises inflation by four-tenths of one percent (0.40). We note that the root mean squared error (RMSQE) from this estimated nonlinear relationship is about 20 percent lower than the RMSQE from estimating a completely backward-looking standard linear Phillips curve over the same sample. In other words, over this sample period an "accelerationist" Phillips curve specification does not perform especially well.

We now use the estimated relationship (2.1) to fit current inflation developments in the United States. Figure 3 plots actual core PCE inflation and predicted inflation from 2008:Q4 onward. To measure π^{LR}, we use the Hoey-Philadelphia survey's median 10-year inflation expectations.[8] The figure shows that while the simple estimated relationship (2.1) fails to capture some short-term inflation dynamics, it does capture the magnitude of the decline in core PCE inflation witnessed since the end of 2008. The CBO estimate of the output gap indicates that resource slack peaked over the course of 2009. Since then, slack has been decreasing modestly. Therefore, according to the nonlinear Phillips curve (2.1), inflation bottomed out over the course of 2010 at a value slightly below 1 percent. With some improvement in the output gap, inflation is expected to revert back to its long-run level. Given the estimated nonlinearity, the reversion is extremely gradual. Core PCE inflation is expected to stay below 1.3 percent as long as the output gap remains below –3 percent.

The projected path for inflation according to (2.1) over the course of 2008–2011 is remarkably close to the range of estimates provided by Stock and Watson (2010), who examine the relationship between inflation and unemployment by looking exclusively at recession episodes over the post-1960 sample. In addition to using a different sample period, resource slack is defined differently and enters the specification linearly. Still, there are some more fundamental similarities between the approach we employ and the one used by Stock and Watson. Most notably, inflation in both exercises is a function of long-run (trend) inflation and an activity gap. Whether this is an accurate reduced-form representation of the inflation process is an issue that will be considered more closely in the next section.

While it may seem far-fetched to project the near-term evolution of inflation based on a Phillips curve relationship estimated over the period 1954 to 1963, we note that recent movements in inflation fall well within the range experienced during this earlier period. Figure 4 depicts the

[8] This series is available from the Board of Governors of the Federal Reserve System in the context of the FRB/US model. The series adjusts the Survey of Professional Forecasters 10-year measure of inflation expectations to take into account the differential between PCE and CPI inflation. The mnemonics for this series in FRB/US is "PTR."

unemployment rate vis-à-vis the four-quarter-ahead inflation rate, expressed as a deviation from long-term inflation. The chart superimposes observations from the period 1954:Q1 to 1963:Q4 with observations spanning the period 2003:Q1 to present. As before, trend inflation in the earlier period is given by the average inflation rate prevailing over that sample. For the most recent period, inflation is measured by the core PCE index, and trend inflation is given by the Hoey-Philadelphia survey measure of 10-year inflation expectations. Merging these two periods yields the following estimate of the nonlinear Phillips curve:

$$\ln(1+0.01*(\pi_{t+4}^{4}-\pi_{t}^{LR})) = \underset{(0.0060)}{0.0386} - \underset{(0.0036)}{0.0231}*\ln(5.3-0.50*\tilde{y}_t), \quad n=69,\ R^2=0.50, \tag{2.2}$$

where standard errors for the estimates are in parentheses. These estimates are very similar to the ones obtained using only the earlier sample. The slope of the curve is somewhat shallower, but the implied near-term evolution of inflation is not materially different from the path depicted in figure 3.

When compared to simple benchmarks, the estimated nonlinear relationship is not devoid of empirical content. Using the estimates in (2.2), the RMSQE of the forecast for inflation over the period 2003:Q1 to present is 33 percent smaller than the RMSQE from a naïve forecast à la Atkeson and Ohanian (2001). In Atkeson and Ohanian's setup, the forecast for four-quarter-ahead inflation is given by the current value of four-quarter inflation. The RMSQE from the nonlinear specification is also 40 percent smaller than the RMSQE from forecasting four-quarter-inflation by just positing an expected value equal to the current level of long-run inflation expectations.[9]

The modest decline in U.S. inflation implied by the simple nonlinear Phillips curve given the current large amount of slack in the economy is the result of two main features of the model. First, the process for inflation is anchored by trend inflation (or long-run inflation expectations). In this respect, the nonlinear specification generates dynamics that are similar to purely forward-looking New Keynesian Phillips curve models (see Fuhrer and Olivei 2010). There are other specifications of the inflation process that do find some support in the data and which entail a less benign inflation outlook, as the next section will show. Even within the confines of this simple setup, the stability of long-term inflation is a crucial factor in limiting the decline in inflation, as no feedback is assumed from the current low readings of inflation to long-run expected inflation. If long-run expectations

[9] Williams (2006) notes that since the mid-1990s core inflation has remained remarkably close to a stable long-run target, and that inflation forecasts based on this long-run target beat random-walk forecasts à la Atkeson and Ohanian.

were to decline, then the implied decline in inflation would be larger. In this regard, it is important to note that during the period from 1954 to 1963, the recessionary episodes were fairly short-lived. For example, in the recession of 1957–1958, the output gap was closed in the first half of 1959. With relatively short downturns, it may not be unreasonable to assume little change in long-run expected inflation, other things equal. But in a situation like the current one in which slack in resource utilization may persist for a long time, such an assumption may prove inaccurate.

The other feature at play in generating a small projected additional decline in inflation is the nonlinear tradeoff between inflation and resource slack. In his original article, Phillips argued that when the unemployment rate is low and labor demand is high, firms will "bid wage rates up quite rapidly." Conversely, Phillips conjectured that with high resource slack, wages will fall "only very slowly" because workers are "reluctant to offer their services at less than the prevailing rates."

The notion of an asymmetric response of wage changes to shocks in the economy was exploited by Akerlof, Dickens, and Perry (1996) to derive time-series implications for inflation. One important difference in Akerlof, Dickens, and Perry's model compared to a specification featuring a nonlinear relationship between inflation and the output gap is that in their model high resources slack is still compatible with rapid disinflation whenever the starting point for inflation is elevated. With high inflation, downward nominal wage rigidity is in fact not binding. When the starting point for inflation, instead, is low as in the current situation, their model has implications for inflation similar to the nonlinear specification considered in this section, as downward nominal wage rigidity limits firms' ability to cut prices.[10] While this feature of Akerlof, Dickens, and Perry's model may be desirable, we note here that the nonlinearity in the relationship between inflation and the output gap which appears to be present in the 1954 to 1963 period may have been at work in other circumstances as well. Figure 5 shows a scatterplot of the CBO estimate of the output gap vis-à-vis the four-quarter-ahead core PCE inflation rate less the 10-year inflation expectations. The figure merges the periods 1964:Q1 to 1969:Q4 and 1981:Q1 to 1997:Q4.[11] Figure 6 shows the same type of scatterplot, but over the period 1973:Q1 to 1980:Q4. Both figures, with the important small-

[10] In order to generate a projected path for inflation as the one depicted here in figure 3, Akerlof, Dickens, and Perry's model, too, needs some anchoring of inflation expectations to a long-run target. Absent such anchoring, their model would predict a mild deflation at levels of the unemployment rate persistently above 7.5 percent (see their figure 3, page 32).

[11] The 10-year expectation measure was not available for the earlier period, and it is constructed by fitting the estimated evolution of the 10-year Hoey-Philadelphia survey measure over the period 1981 to 2006 to that earlier sample. The estimated process for the series is highly persistent, with some additional influence coming from actual core inflation, the monetary policy stance, and the output gap. This series is constructed by the Board of Governors of the Federal Reserve System. See footnote 9.

sample caveat we already mentioned, are suggestive of a sacrifice ratio increasing when the output gap is negative. For this type of relationship to emerge, it is crucial to consider the deviation of inflation from long-run expected inflation. Through this lens, recessions are periods when inflation reliably falls, but the change in the inflation gap when output is below potential appears to be smaller than the change in the inflation gap when output is above potential. Still, it is important to stress that over the periods represented in figures 5 and 6, long-run inflation expectations were moving as well. The change in the inflation gap may well be muted with high resource slack, but if long-run inflation expectations are moving considerably so will actual inflation. The other implication is that for these periods it could prove more appropriate to appeal to some form of real, rather than nominal, wage rigidity in order to justify a nonlinear relationship between inflation and real activity.

3. Inflation in a Low-Inflation Environment: The Japanese Example

As discussed above, over the past 50 years the United States has had relatively little experience with an annual inflation rate below 2 percent. Consequently, it is difficult to garner much evidence that would help policymakers gauge whether the inflation rate may behave differently as it approaches zero. However, Japan provides a potentially useful recent example of a developed country that has experienced a prolonged recession and an extended period of very low (indeed negative) inflation. While one must be cautious about drawing parallels too closely between the United States and Japan, the Japanese experience may provide some clues as to how U.S. inflation might evolve in the years to come.

We begin by examining key macroeconomic data for Japan over the past 30 to 40 years. Figure 7 displays some of the data central to our analysis. The top panel shows both headline and core CPI inflation for Japan, where "core" is defined as all consumer items excluding food and energy, along with a measure of the output gap. Estimates of Japan's output gap are hard to come by. Some estimates suggest that output returned above potential in the mid-to-late 1990s, although this seems implausible. The output gap presented here was kindly provided by Sachihiro Hayashi at the Cabinet Secretariat in Japan. Unlike most other published estimates, it suggests what we consider a plausible history for the Japanese output gap—that is, a gap that implies protracted periods in which output falls significantly below potential—and performs reasonably well in an array of estimates presented below.

In some of the Phillips curves that we consider, we also use estimates of real marginal cost (proxied by the labor share of income, as is conventional in the Phillips curve literature), the relative price of imported energy goods, and the exchange rate. These data are readily available and obtained from standard sources.[12] Of more interest and less readily available are measures of inflation expectations. The bottom panel of figure 7 displays both short-run (one year) and long-run (six to 10 years) inflation forecasts, taken from the Consensus Forecast database.[13] These data are available only from the second half of 1989, and are collected semi-annually. Other horizons (2–5 years) are included in the dataset, as are forecasts for a limited set of other macro variables. We use the output growth forecasts for forecast horizons of 1–5 years to create estimates of the expected output gap.

Note immediately a few features of the data:

1. Over the 1998 to 2010 period, inflation has clearly fallen significantly and persistently below zero.
2. This decline in inflation occurred despite long-run inflation expectations that have remained around 1 percent for the past 15 years.
3. However, inflation did not spiral downward, despite the presence of an (estimated) negative output gap throughout most of this period.
4. While the 6-10 year forecast has been wrong (on the optimistic side) for many years, the one-year forecast tracks actual inflation reasonably well.

<u>Reduced-Form Models: Old-Style Phillips Curves and the Phillips Correlation</u>

Consider a standard backward-looking Phillips curve, where inflation depends on lagged inflation, usually with a coefficient of one, and on the output gap, thus:

$$\pi_t = \pi_{t-1} + a\tilde{y}_t \tag{3.1}$$

Lagged inflation serves as a proxy for expected inflation, or for frictions in the price-setting mechanism. The logic of this specification implies that inflation will continue to rise (fall) as long as the output is positive (negative). This dynamic has been dubbed the "accelerationist" Phillips curve, as it posits a relationship between the *change* in inflation (and thus the second-difference or acceleration in the price level) and the gap. Examining the top panel of figure 7, it would seem

[12] See the data appendix for details.
[13] Note that in the bottom panel of the figure, the forecasts are displayed on the date the forecast is made.

difficult to reconcile such a specification with Japanese inflation experience of the past 30 years. Despite a persistently negative output gap since the early 1990s, inflation appears to have been bounded below by –2 percent.

More formally, standard backward-looking Phillips curves show a marked deterioration in fit in recent decades. Table 2 provides a simple illustration of this point. The first set of columns displays the estimated backward-looking Phillips curve for Japanese core inflation from 1971 to 1989. As the estimates suggest, the unit sum restriction on lagged inflation (three quarterly lags) is not rejected, and the output gap displays a sizable and significant coefficient sum. The second set of columns show the estimated shift in these parameters from 1990 to the present. The effect of the output gap is essentially eliminated, with great precision.[14] The table suggests other significant shifts as well, but the change in the estimated effect of the output gap is economically quite important, and we consider it in more detail below.[15]

Figure 8 displays the in-sample and out-of-sample fit for the same Phillips curve over a variety of subsamples from 1980 to the present. The Phillips curve includes three lags of core inflation, the change in the relative price of imported energy goods, and lags of the estimate of the output gap displayed in figure 7. As the figure indicates, forecasts made employing a Phillips curve estimated through the mid-late-1990s consistently and significantly underforecast the level of inflation in ensuing years. The implication is that the persistent negative output gap would have implied more pronounced declines in inflation, but these evidently did not occur. Note that the trajectory of the inflation forecast is consistent with a continued downward spiral—as long as the output gap remains negative, inflation continues to fall in the forecast.

Why does Japan's Phillips curve miss so badly during the more recent decades? As suggested by the shift test above, the basic accelerationist nature of the relationship is not present anymore. The recent data might suggest a purely forward-looking rational expectations model of inflation in which inflation depends on the average (discounted) future output gaps, so that as the level of the expected output gap increases, inflation would rise. This possibility is examined in more detail below.

[14] A flattening of the Phillips curve in Japan since the 1990s is also observed by De Veirman (2009).

[15] At some econometric risk, we splice the OECD's estimated output gap for 1970-1979 with the Cabinet Office's gap from 1980 to the present. We allow for shifts in the pattern of lagged inflation coefficients as well, but these are not individually significant. They are constrained to sum to zero to preserve the unit sum constraint on lagged inflation. Note that unknown breakpoint tests of the Japanese Phillips curve suggest a break in the standard relationship in the late 1970s, well before the asset price collapse in the late 1980s that preceded Japan's great recession.

Models that Include Survey Expectations

The concept of anchored expectations plays an important role in discussions of the expected path of U.S. inflation. As highlighted in "trend inflation" models such as those proposed by Cogley and Sbordone (2008), sufficiently anchored inflation expectations can act as an offset to the downward pull from a significant output gap and/or declining marginal costs. The following subsections examine the role of expectations in the inflation process in Japan, using both surveys of professional forecasters and rational expectations models as alternative means of measuring expectations. The goal is to derive implications from the Japanese experience that may cast light on the future course of U.S. inflation.

Returning to figure 7, the bottom panel suggests that long-run inflation expectations (the red line) have remained well-anchored in Japan, fluctuating relatively little around their average of 1.2 percent over the past 15 years (the average is shown in the dashed red line). This observation raises the question of how long-run inflation expectations are formed. It is difficult to arrive at a fully satisfactory explanation, but one hypothesis is that Japanese forecasters place considerable faith in the Bank of Japan to ultimately deliver positive inflation—more faith than might be warranted given the constraints under which it labors. Since the inception of the survey in 1989, long-run expectations have always foreseen positive inflation in the 6–10 years ahead. That forecast has had the wrong sign for most of the past 15 years, yet even the most recent reading, in the aftermath of a prominent drop in Japanese output, expects inflation to recover to about 1 percent on average six to 10 years from now.

While the accuracy of the long-run Japanese forecast is of interest, of more relevance to this paper is the role that these expectations play in influencing realized inflation. To be sure, anchored long-run expectations have not prevented Japanese inflation from falling below zero and remaining there for the last dozen years. But as already discussed before, Japanese inflation has not *continued* to decline, and this observation raises the possibility that anchored long-run expectations have kept inflation from dropping further than they might have in the presence of a large output gap.

To examine the influence on realized inflation of both short- and long-run expectations, we estimate simple Phillips curves that explore the relationship among core inflation, expectations, the output gap, marginal cost, and important relative price shifts.[16] The Phillips curves take the form

$$\begin{aligned} \pi_t &= a\pi_t^{1y} + b\pi_t^{LR} + c\pi_{t-1} + d\tilde{y}_{t-1} + es_t + fx_t \\ a+b+c &= 1, \end{aligned} \tag{3.2}$$

where s_t denotes marginal costs, $\tilde{y}_t$ represents an estimate of the output gap, and x_t represents relative price variables (energy prices and the exchange rate). We entertain both the output gap and marginal costs as driving processes, as the literature is somewhat unsettled on this matter. The sum of the coefficients on survey expectations and lagged inflation are constrained to sum to one, although we test the significance of that constraint throughout. In preliminary tests, we find that marginal costs often enter these regressions significantly; the output gap never does, so we exclude it from the results presented below. In addition, the relative price variables are rarely significant, so we exclude these from the analysis. We present both full-sample (table 3) and rolling regression estimates (figure 9).

Note that as highlighted above, now-common specifications for inflation typically express inflation relative to "trend inflation," see Cogley and Sbordone (2008). If inflation exhibits a trend component, perhaps attributable to a time-varying inflation goal, then it is appropriate to write the inflation model so that all inflation terms are expressed as deviations from the trend term. In the context of these datasets, long-run inflation expectations could serve as a reasonable proxy for trend inflation.

While the theory supporting such specifications is compelling, the data for this sample is much less so. Consider a trend-inflation version of the equation above:

$$\pi_t - \pi_t^{LR} = a(\pi_t^{1y} - \pi_t^{LR}) + (1-a)(\pi_{t-1} - \pi_{t-1}^{LR}) + es_t. \tag{3.3}$$

This equation implies restrictions on the way in which long-run expectations enter: Re-arranging this equation, it can be shown that the coefficients on long-run inflation in periods t and $t-1$ should be $(1-a)$ and $-(1-a)$ respectively. A test of this restriction fails to reject it, yielding a *p*-value of 0.11. However, the coefficients for long-run inflation expectations enter insignificantly, both individually and jointly, likely explaining the inability to reject the opposite-sign restriction implied by the trend

[16] The working paper version of this paper provides simple VAR analysis of the covariation among these key variables. The results are consistent with the estimates presented here.

inflation model. [17] Altogether, it seems empirically well-justified to abstract from the issues implied by trend inflation for Japanese Phillips curves over this period. The results that follow employ the simpler specification, that is equation (3.2) with $b = 0$ and $d = 0$.

The estimates in the top panel of table 3, along with rolling regression results (not shown), suggest that long-run expectations play little or no role in the evolution of Japanese core inflation.[18] For this reason, we report estimates that exclude long-run inflation expectations. Estimation results for the full sample are presented in the bottom panel of table 3, and rolling regression results are shown in figure 9.

Our findings suggest a strong role for the one-year expectation, both for the full sample and for the subsamples covered by the rolling-regressions of figure 9. The *p*-value (not shown in the figure) for the one-year expectation is always near zero. Lagged inflation plays a small but moderately significant role throughout most of the sample. For the last dozen years, the role of lagged inflation becomes insignificant, the marginal costs variable gains significance, and the *p*-value of the unit sum restriction drops toward zero, suggesting this restriction is violated in the most recent decade. The finding of a strong role for the one-year expectation is robust to concerns about multicollinearity between the output gap and the one-year expectation. In the bottom panel of table 3, substituting the output gap for the one-year expectation results in an insignificant coefficient on the output gap, and a decline in the R^2 from 0.79 to 0.39. The one-year expectation appears to capture an important and independent effect on inflation.

While the backward-looking accelerationist Phillips curves of the preceding section are poor predictors for the Japanese inflation experience of the past 20 years, the survey expectations models that use real marginal costs as a driving variable achieve more success. For this model, figure 10 displays the in-sample fit for the full-sample estimates reported in table 3, and the out-of-sample fit for the last five years of the sample. The out-of-sample fit suggests the model is reasonably stable (given the limited number of semi-annual observations available to test this hypothesis), and certainly far more stable than the Phillips curve results presented in figure 8.

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[17] The relative stability of the 6–10-year inflation expectation over the past 20 years also suggests that this variable will add little to the model. Equations estimated in the deviation form (3.3) produce results that are very similar to those in tables 4 and 5. Specifically, in table 3 the estimated coefficients on the one-year expectation and lagged inflation are 0.92 and 0.08 respectively; the coefficient on marginal costs is 0.16. In table 5, the estimated effect of the rational expectations term is once again zero, and the weights on the one-year expectation and lagged inflation are 0.7 and 0.3, respectively. The marginal costs estimate, at 0.22, is a bit higher than the OLS estimate.

[18] The working paper version of this paper contains a more complete set of results, including the rolling-regression estimates of equation (3.2) referenced in the text.

Because the success of this model rests heavily on the inclusion of the one-year survey expectation of inflation, it is important to better understand what information is incorporated in the one-year survey. Reduced-form regression models of the one-year expectation reveal that one-year inflation expectations are well-explained by the following regression equation:

$$\pi_t^{1y} = a\pi_{t-1}^{1y} + b\pi_{t-1} + cs_t + d\tilde{y}_t + c_0 . \tag{3.4}$$

The estimated equation using semi-annual data from 1990:H1 to 2009:H2 is (HAC standard errors in parentheses)

$$\begin{array}{ccccccccccc} \pi_t^{1y} & = & 0.31\pi_{t-1}^{1y} & + & 0.18\pi_{t-1} & + & 0.056s_t & + & 0.20\tilde{y}_t & + & 0.55 \\ & & (0.13) & & (0.11) & & (0.039) & & (0.035) & & (0.11) \end{array}$$

$$R^2 = 0.88.$$

The one-year expectation is correlated with lagged inflation, real marginal costs and the current output gap, and exhibits very modest persistence, as evidenced by the significant but small coefficient on the lagged dependent variable. The in-sample fit of this equation, which is quite tight, is displayed in the bottom panel of figure 10. Note that inclusion of the 6–10-year inflation expectation adds marginally to the fit of the equation in the first five years of the sample. For the period since 1995, the coefficient for the long-run expectation is estimated very imprecisely, and adds nothing to the fit of the one-year equation. Once again, we exclude long-run inflation expectations for the balance of this section.

The implication of this simple depiction of one-year Japanese inflation expectations is that they need not be well-tied the long-run expectations of inflation as measured by the Consensus forecast survey. Expectations respond to realized inflation and to the output gap with a lag. As the output gap improves, inflation improves, but with a lag, given the dependence on its own lag and on realized inflation. How well such an expectations mechanism, coupled with the inflation equation estimated above, can explain the Japanese inflation experience in a more fully fleshed-out model remains to be seen.

Structural Models

While the old-style backward-looking Phillips curve fits the Japanese inflation data quite poorly, a model with survey expectations achieves some success. How well would the current generation of forward-looking models explain Japanese inflation behavior?

We construct a simple rational expectations model that encompasses both the NKPC model—with and without indexation—and a survey expectations-based model. An interesting aspect of this exercise is that it allows us to examine the extent to which the survey expectations mimic the expectations implied by the rational expectations NKPC model for Japanese data. The inflation equation is

(3.5) $$\pi_t = aE_t\pi_{t+1} + b\pi_t^{1y} + (1-a-b)\pi_{t-1} + cs_t .$$

If $a=1$ and $b=0$, then the rational expectations NKPC is a reasonable approximation for Japanese inflation data. If $a=0$ and $b=1$, then the model with one-year survey expectations serves as a better depiction of Japanese inflation. With $a=0$ and $b=0$ the model depends only on lagged inflation, as in an old-style Phillips curve model. A host of intermediate combinations are of course possible.[19, 20] Note that we have reverted to marginal costs as the driving variable in the Phillips curve. This is largely a matter of empirical fit—initial estimates suggest that the output gap enters insignificantly and often with the wrong sign in this model. Thus we use model-consistent forecasts of marginal costs, in contrast to the previous exercise, which uses Consensus survey forecasts of the future values of the driving variable.

Since we need to generate model-consistent expectations of inflation for period t+1 and we need to be concerned about the possible simultaneous determination of marginal cost (and potentially of the survey variable) and inflation, we close the model with the reduced-form equation for one-year inflation expectations as described above, and with VAR equations for marginal costs and the output gap estimated over the full sample (1990 to 2009). Thus the full model that we employ for estimation, with error terms suppressed, is:

[19] Using somewhat different models, samples, and methods, Roberts (1997) examines the role of survey expectations in forward-looking Phillips curves for the United States.

[20] In preliminary estimation exercises, the lag of *total* inflation, rather than core, enters significantly in the inflation equation. Thus we allow inflation to respond to lagged total inflation, and link total inflation to core inflation via an estimated error-correction equation.

$$
\begin{aligned}
&\pi_t = aE_t\pi_{t+1} + b\pi_t^{1y} + (1-a-b)\pi_{t-1}^T + cs_t + c_0 \\
&\pi_t^{1y} = d\pi_{t-1}^{1y} + e\pi_t + f\pi_{t-1} + g\tilde{y}_t + g_0 \\
&\pi_t^T - \pi_{t-1}^T = j(\pi_{t-1}^T - \pi_{t-1}) + k(\pi_{t-1} - \pi_{t-2}) \\
&s_t = \sum_{i=1}^{2} \mathrm{B}_i X_{t-i} \\
&\tilde{y}_t = \sum_{i=1}^{2} \Gamma_i X_{t-i}\ ,
\end{aligned}
\tag{3.6}
$$

where π_{t-1}^T is the lagged value of overall inflation (which performs better than lagged core inflation in initial regressions), and X is the vector of variables in the two-lag VAR (inflation, output gap, marginal costs). The driving variable for the one-year expectation is the output gap, as initial estimates find an insignificant role for marginal costs. This result is somewhat puzzling, as we find the opposite results for the Phillips curve. The reduced form inflation equation estimated above suggested some role for both marginal cost and the output gap; this system parses those roles into effects on inflation and inflation expectations respectively.

We estimate the VAR coefficient matrices Β and Γ jointly with the other parameters via maximum likelihood.[21] The results for the first three equations in (3.6) are presented in table 4.[22] The estimates find the weight on the forward-looking (model-consistent or rational expectations) component to be small and imprecisely estimated. There appears to be a small role for indexation or lagged inflation, while two-thirds of the weight on the expectations terms falls on the one-year expectation. This result mirrors the results in the reduced-form equations presented above. The contribution of the short-term expectations is again significant: The likelihood ratio test for the hypothesis that $b = 0$ is rejected with a *p*-value of 0.02.[23]

For some purposes, it would be reasonable to collapse the results for the one-year expectation into the single-equation Phillips curve of equation (3.4). The significant estimated coefficient for the one-year expectation would imply an augmented role in the Phillips curve for

[21] The working paper version contains estimates in which we hold the VAR parameters at their OLS estimated values. There is no qualitative difference in the results from the joint estimation. The standard errors for the VAR parameters and some of the structural parameters are larger, but the results are otherwise the same. This result is not surprising given the lack of feedback from the structural equations into the VAR equations.

[22] Nunes (2010) also explores the role of survey expectations in New Keynesian Phillips curves. His results differ somewhat from ours; for his GMM estimates, this may be due to weak instruments (see Fuhrer 2011).

[23] The specification appears well-behaved in most dimensions. For example, the estimation residuals are approximately white noise, with Q(12) *p*-values for the five errors of 0.60, 0.72, 0.89, 0.61 and 0.51 respectively. The in-sample fits are all tight. As noted above, a version of this model that expresses the inflation variables as deviations from long-run expectations does not differ qualitatively from the version presented in table 5.

lagged inflation and the output gap.[24] But this is just the equivalent in a rational expectations framework of solving for the model-consistent expectations in terms of observables to arrive at a single-equation reduced-form representation for inflation that is consistent with the restricted structural model. While one could then use the reduced-form inflation equation as a forecasting vehicle, the point of the modeling in this paper is to understand the way in which expectations—survey -based or model-consistent—enter structural Phillips curves, not to arrive at a better reduced-form forecasting equation for inflation.[25]

In sum, Japanese inflation appears to be best modeled not as a backward-looking accelerationist Phillips curve, nor as a forward-looking NKPC, nor as a hybrid NKPC. Instead, inflation depends on the one-year survey expectations, which evolve according to the dynamics described above, and do not replicate either old-style or NKPC-style Phillips curves.

4. How Much Does the United States Look Like Japan?

Up to this point, our inflation modeling has been suggestive, in the sense that an economically developed country like Japan clearly can experience a sustained bout of deflation despite long-run expectations that are well-anchored above zero. More important to the evolution of inflation is the short-run (one-year) expectation, which in Japan appears to be driven by recent realizations of inflation and the output gap.

How well would such a model fit the United States in recent years? Surprisingly, the answer appears to be quite well. While we will explore the dynamics of U.S. inflation and inflation expectations in more detail shortly, the spirit of the first exercise in this section is to simply apply the Japanese inflation model to U.S. data to see how well the model explains recent history. Panel *A* of table 5 presents estimates of a Phillips curve that closely parallels the estimated model in table 3 for Japan.[26] Note that in initial regressions, as in the Japanese dataset, long-run (10-year average) expectations from the SPF survey enter with a very small coefficient (0.01), which is imprecisely estimated (standard error 0.25), also paralleling the evidence from Japan. Thus we exclude long-run inflation expectations from the regressions in table 5. The issues that arise from considering so-called trend inflation models are present here as they were in the case of Japan. The data for the

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[24] The positive coefficient on lagged expected inflation implies additional lags of all the variables in the Phillips curve as well, although this lagged coefficient is relatively small.

[25] In a rational expectations framework, one can also solve for the single-equation reduced form for inflation, expressed in terms of observables.

[26] See the data appendix for details on data construction and sources.

United States also fail to reject the restriction implied by the trend inflation model, with a *p*-value of 0.15. But like Japan, the long-run (10-year) expectations are estimated very imprecisely, and are far from significant either individually or jointly. As a consequence, we exclude 10-year inflation expectations from the equation determining inflation for the United States. The estimates that follow employ the simpler, non-trend-inflation specification outlined above. The one-year expectation takes on a similar coefficient of about two-thirds. Excluding the one-year expectation from the regression decreases the R^2 from 0.69 to 0.57. The one-year expectation clearly explains a quantitatively significant and statistically independent source of inflation variation.[27]

Panel *A* of table 5 also displays estimates for the one-year expectation for U.S. inflation, which is pivotal in determining U.S. inflation, just as it is for Japan. The parallels are striking. In the core inflation equation, the coefficients on the one-year (SPF) survey expectation and lagged inflation are extremely close to those estimated for Japan. The coefficient on marginal costs is smaller, but still likely plays an important role. The one-year expectation similarly depends on lags of realized inflation and the output gap.[28] Figure 11 displays the fitted values for both of these regressions, with the post-2006 out-of-sample fit for the core inflation regression depicted in the top chart.[29] There is no obvious sign of out-of-sample degradation.

In parallel with the exercises conducted for Japan, we estimate a version of the model for U.S. data that allows the rational expectation of next period's inflation to enter the inflation equation. The model parallels the Japanese model in the previous section, and is displayed in panel *B* of table 5. Broadly speaking, the results are quite similar to the OLS estimates above, which is not surprising given the absence of a rational expectations effect in the equation.[30] Inflation is well-

[27] A version of the model in which we express inflation as the deviation from trend inflation—in this case, as a deviation from the SPF 10-year inflation expectation—produces results that are qualitatively similar. Specifically, the coefficients on the one-year lagged inflation and marginal cost are 0.73, 0.27, and 0.057, remarkably close to the estimates in table 6. The relative stability of the SPF 10-year expectation suggests that this adjustment is close to subtracting a constant from the inflation measures, which of course should not affect the results much.

[28] One can improve the fit of the equation by including the lagged one-year expectation, but the estimated coefficient in this case is about 0.7, which seems like an undue reliance on lagged expectations.

[29] The estimates in panel *A* of table 5 are OLS estimates. GMM estimates of the same model yield very similar results, with the sum of coefficients on the inflation and output gap terms insignificantly different from the OLS estimates. The *J*-statistic fails to reject the over-identifying restrictions. Full information estimates appear in panel *B* of the table. The out-of-sample fit is generated by estimating the regression from 1990 to 2006, and then dynamically simulating the model with these estimated coefficients for 2007–2010:H1.

[30] Note, however, that the maximum likelihood technique employed in these estimates takes full account of the simultaneity implied by the presence of the contemporaneous inflation and output terms in the equation for one-year inflation expectations. The estimates for *a, b* and *c* are constrained to be greater than or equal to zero, as it is difficult to develop a structural interpretation in which any of these expectations proxies would enter negatively. Estimates for the model in which inflation variables are expressed as deviations from the 10-year expectation deliver remarkably similar results. The estimates for *a, b* and *c* are 0.00, 0.78 and 0.52 respectively. In the one-year expectation equation, the sum of

modeled as depending on the one-year survey expectation—with literally no role for the rational expectation that is consistent with this model—and a modest dependence on lagged inflation that may reflect indexation or rule-of-thumb price-setting. The likelihood-ratio test for the restriction that $b = 0$, implying that other factors can equally replicate the explanatory power of the one-year expectations, develop a *p*-value of 0.000. The reduced-form evolution of the one-year inflation expectation is quite similar to that of Japan, in that one-year expectations respond sluggishly to realized inflation and the output gap. It is of some interest that for the United States, the change in the output gap may be as important as the level of the output gap in influencing one-year expectations, as evidenced by the small sum of the lag coefficients, but very high joint significance of the output lags.

While this specification is essentially transplanted from the Japanese example, using the sample for which expectations data are available in Japan, the model actually fits U.S. data from 1981–1990 quite well. Figure A1 in the appendix displays two tests of the fit prior to 1990. The top panel displays the in-sample fit, re-estimating the equation in the top panel of table 5 from 1981–1989; the bottom panel displays the fit of the equation estimated from 1990–2010 on the sample 1981–1989. In both cases, the model fits the pre-1990 data reasonably well, although the model misses significant fluctuations in inflation during the 1982–1984 disinflation.

It is striking that long-run inflation expectations explain little or nothing of recent U.S. inflation fluctuations, once one-year expectations are taken into account, contrary to the prevailing view that it is essential for the stability of inflation that long-run expectations remain well-anchored. However, we find evidence that long-run expectations may have an *indirect* effect on inflation, via their effect on one-year expectations. This finding stands in contrast to that for Japan, in which the long-run expectations exhibit little relationship to either realized inflation or to short-run expectations of inflation.

Table 6 presents the results from re-estimating the equation for one-year inflation expectations, allowing for the influence of long-run expectations.[31] Long-run expectations clearly play a role in determining one-year inflation expectations for the United States, even though these never enter significantly in any of the estimated inflation equations. The R^2 for the same regression excluding long-run inflation expectations declines from 0.92 to 0.84. Now the estimated sum of the

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the coefficients on realized inflation is smaller, and the sum of the coefficients on output larger than in the estimates presented in table 5.

[31] The authors thank Peter Hooper for pointing out this correlation during his discussion of the paper at the Boston Fed's annual economic conference held in October 2010.

output gap coefficients is larger and significantly different from zero, indicating a role for the level of the output gap, as in the Japanese data. The fitted value for the equation, displayed in the bottom panel of figure 11, more closely adheres to the actual one-year expected inflation rate than does the Japanese-style model.

In earlier work (Fuhrer and Olivei 2010), we show that long-run SPF inflation expectations are best modeled as a very slow moving average of realized inflation. Thus the presence of these expectations in the estimated one-year expectation equation above represents two important influences on one-year expectations: (1) the influence of a slow moving average of past inflation, as reflected in the results in table 6, and (2) the tendency for one-year expectations to gradually revert towards long-run expectations, which presumably reflects the central bank's inflation goal. Note that in the past decade, the 10-year expectation has exhibited remarkably low variance, remaining extremely close to a steady 2.5 percent throughout (since 2000:Q1, the series has displayed a mean of 2.5 and a standard deviation of 0.055). Thus in recent years, the 10-year expectation largely represents the forecasters' assessment of the central bank's stable inflation goal.

To recap, the dynamics of inflation for both the United States and Japan may be summarized as follows:

1. Conditional on the assumption that the error term is white noise, inflation is not well-modeled by an old-style accelerationist Phillips curve, or by a forward-looking, rational expectations New Keynesian Phillips curve.
2. Inflation may be sensibly modeled as depending on the one-year (survey) expectation of inflation with a weight of between 2/3 and 3/4, and lagged inflation with a weight of between 1/3 and 1/4.
3. Inflation also depends on marginal costs.
4. The one-year expectation responds sluggishly to current and lagged realized inflation rates, and to current and lagged output gaps. In the United States, the one-year expectation appears to revert gradually to the long-run expectation, which in recent years has shown remarkable stability.

5. Implications for the United States in Coming Years

Of interest is what the implications of such a model are for the evolution of inflation in the United States given the current circumstances of a very low-inflation environment where monetary policy is constrained by the zero lower bound. It is conceivable that slowly moving one-year expectations, combined with a large and persistent output gap and depressed marginal costs, could lead to a very slow adjustment of inflation towards its long-run target. We next simulate a model that incorporates these dynamic properties of inflation, augmented by a conventional monetary policy rule in which the policy rate is constrained not to fall below zero.

The model comprises the inflation and expectation equations in tables 5 and 6 above, plus a conventional policy rule that imposes the zero lower bound on the short-term policy rate r_t [32]

$$r_t = \max\left\{[a_\pi(\pi_t - \bar{\pi}) + a_y\tilde{y}_t + \bar{r}], 0\right\}, \tag{5.1}$$

where $\bar{\pi}$ is the central bank's (assumed constant) inflation goal, and the parameters of the policy rule are set to $a_\pi = 2, a_y = 1.0$, which is a bit more aggressive than the canonical Taylor (1993) parameters but is consistent with empirical estimates that include more recent years. The inflation goal is set to 2 percent, and we take long-run inflation expectations to be equal to the inflation goal. The equilibrium real rate is 2 percent, so that the equilibrium nominal policy rate is 4 percent. The model is closed with a "hybrid" I-S curve of the form

$$\begin{aligned} \tilde{y}_t &= \omega_b\tilde{y}_{t-1} + \omega_f E_t\tilde{y}_{t+1} - \sigma(r_t - E_t\pi_{t+1} - \bar{\rho}) + u_t \\ u_t &= \rho_y u_{t-1} + \eta_t \,. \end{aligned} \tag{5.2}$$

Here $\bar{\rho}$ is the equilibrium real rate of interest, and σ is set to 0.05, consistent with many estimates in the dynamic stochastic general equilibrium literature. Note that in introducing a policy rule in order to enforce the zero lower bound, we must eliminate the constants from the regression specifications for inflation and expected inflation in table 5 above. The equilibrium levels of inflation and expected inflation are determined by the inflation target in the policy rule.[33]

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[32] The zero lower bound is imposed by using the Heaviside function H(x) = 1 if x >= 0, = ε if x < 0. Because the solution procedure utilizes analytic derivatives, it is convenient to use a function with a well-defined derivative. The derivative of the Heaviside function is the Dirac function.

[33] In addition, we impose the constraint that the coefficients on lagged inflation in the expectation equation sum to one; this ensures that the model converges to a reasonable steady-state. Absent this restriction, the steady-states for inflation and inflation expectations can differ, and the steady-state values for other nominal variables can differ from the inflation goal in the policy rule. While these small differences likely are difficult to detect in the econometric estimates, they clearly matter for the model's steady-state, so we impose them in this simulation. They make only a modest difference in the simulation—inflation still remains below its central bank-dictated goal for more than ten years.

The model is simulated with initial conditions that are intended to mimic macroeconomic conditions at the beginning of 2010, namely that:

- Inflation is 1 percent;
- The federal funds rate is just above zero;
- The output gap is a considerable -7 percentage points;
- Short-term inflation expectations are also 1 percent.

The evolution of the key variables in this simulation depends importantly on the specification of the output equation. If one assumes that the shock *u* is uncorrelated over time, and that output is purely forward-looking, then the output gap jumps immediately above zero, inflation begins to rise, and the policy rate is generally not constrained by its zero lower bound. However, in more realistic depictions of output—either a highly correlated error term or a substantial weight on lagged output, ω_b—the combination of a high dependence on short-term expectations and the somewhat sluggish adjustment of expectations to current and lagged conditions implies the outcome depicted in figure 12.[34]

The outcome of this simulation's predicted path for U.S. inflation is striking as, using a model that fits U.S. inflation experience of the past 20 years quite well, it mimics to a great extent the historical experience of Japan. The model suggests a prolonged period for which inflation is negative and the policy rate is pinned at the zero lower bound as the output gap recovers gradually. At first, inflation falls with one-year inflation expectations, which decline as a result of the large and negative output gap. Over time, as the output gap improves, inflation remains low because of the effect of lagged inflation on the one-year expectation. While inflation and expectations ultimately return to the 2 percent inflation goal (the long-run inflation expectation) in the model, the simulation suggests that the return to long-run equilibrium could take nearly a decade.

Which Case to Believe: The Optimistic or the Pessimistic One?

The results in section 2 suggest a somewhat more optimistic take on the likely evolution of U.S. inflation in the near-term: Due to the nonlinearity of the model and the reversion to "trend

[34] At the time of the initial writing in early 2011, most private forecasters envision a protracted period of elevated unemployment that extends several years into the future. This would normally imply a significant output gap over the same period, even with due consideration to the possibility that structural unemployment has risen and/or potential output has fallen.

inflation," inflation is expected to dip no lower than about ¾ of one percentage point (0.75), returning gradually to 2 percent in ensuing years. The simulation presented in section 5 parallels developments in Japan and suggests a less optimistic outcome, with U.S. inflation dipping well below zero for quite a few years and then only gradually returning to the central bank's target. Which outcome is more plausible?

The two models are not nested, so classical hypothesis testing is not possible. However, simple tests for non-nested hypothesis testing developed by Davidson and MacKinnon (1981) can shed some light on the relationship between the two models. The intuition behind these tests is straightforward: A regression of the dependent variable (y_{1t} in Table 7) on the regressors of one model (X_{1t}), along with the fitted values from the second model ($\hat{F}_{2t}$), allows one to examine the extent to which the second model provides explanatory power not contained in the first. The reverse test, in which the second model's regressors and the first model's fitted values are used to explain the dependent variable, provides an assessment of the extent to which the first model provides explanatory power not contained in the second model. A number of possibilities can arise—one model may "dominate" the other, if only its fitted values enter significantly in the test regressions. Alternatively, it may be that neither model dominates, or that both models add important information to the other.

We run this simple test for the nonlinear model of section 2 and the U.S. inflation model from table 5. The tests are computed for samples from 1991:Q1 to selected end dates over the past seven years, to test the robustness of the results for any one sample. The results are summarized in table 7. For all samples in the table, the hypothesis that the fitted values of the Japanese-style model can be excluded from the nonlinear Phillips curve is rejected with extremely high confidence. The opposite is not true for excluding the fitted values from the nonlinear Phillips curve from the Japanese-style model. The results suggest that the Japanese-style survey expectations model dominates the nonlinear Phillips curve. Sadly, the pessimistic case wins out.

While we should be careful in applying lessons from Japan to the U.S. experience, the Japanese data may serve as a cautionary tale for U.S. policymakers. It is worth noting that while the models examined in this section empirically generate the "inflation-floor" outcomes that characterize Japan's recent economic history, these leave open the fundamental puzzle of why inflation would behave as if it had a floor or lower bound. The next section examines a possible explanation for the presence of such a floor.

6. Wage Rigidity and Its Effects on Inflation

As the previous section highlights, there is mounting evidence that the inflation process may change as the rate of inflation approaches zero. Until the recent experience in the United States, Japan represented the best test case study of this possible effect. Since Japan's real estate collapse in the late 1980s, its inflation rate appears to be affected by short-run inflation expectations and the output gap, but the role of longer-run expectations is difficult to discern. The United States has been flirting with low levels of inflation since the late 1990s, with inflation near zero twice in the past decade, but never (to date) dipping below zero even with large estimated output gaps. In contrast to the Japanese evidence, many economists suggest that the relatively stable behavior of U.S. inflation may be attributed to the anchoring effect of fairly stable long-run inflation expectations. Several hypotheses have been forwarded for the stability of long-run expectations, such as increased Fed credibility, a smaller effect of the output gap on inflation, inattentiveness to inflation on the part of price-setters when inflation is near zero, or just good luck. This section revisits a more traditional explanation: downward nominal wage rigidity.[35]

Speculation about downward nominal wage rigidity has a storied past in macroeconomics, both theoretically and empirically.[36] If present, wage rigidity could become more problematic as inflation falls toward zero. When nominal wages and inflation are expanding robustly, significant reductions in the real wage are possible without declines in the nominal wage. As inflation approaches zero, real productivity-adjusted wages can only fall through nominal wage declines.[37] If nominal wages are rigid downward even in the face of significant output gaps, inflation may become rigid downward. If the stickiness of wages is widely recognized, inflation expectations may also become rigid downward at low rates of inflation. Hence, nominal downward wage rigidity can help produce very different inflation and output dynamics.

At the individual level the evidence suggests, though not uniformly (McLaughlin 1994; Lebow, Saks, and Wilson 2003), that individual wages are rigid downward. From examination of

[35] The attribution of inflation's stability to the supposed anchor provided by long-run inflation expectations (in both Japan and the United States), even as actual inflation remains below that expectation for long stretches of time, is a profession of faith in the anchoring power of the central bank's long-run inflation objective. Interestingly, many adherents to this faith quickly turn into apostates once the inflation rate rises *above* the long-run expectation for any length of time.

[36] The modern reference for a theoretical and empirical application of downward nominal wage rigidity to inflation dynamics is Akerlof, Dickens, and Perry (1996).

[37] With productivity growth, a firm's costs and prices may fall even without cutting nominal wages. In the extreme, the lower bound on inflation would look something like the negative of the growth rate of productivity.

individual wage changes, as in Gottschalk (2004), Card and Hyslop (1997), and Kahn (1997), to employer interviews (Bewley, 1999), the evidence suggests a reluctance of workers to accept a decline in wages. This reluctance can be clearly illustrated by examining the distribution of wage changes of job stayers from the PSID, figure 13. Although there has been significant discussion of measurement error in the PSID, (for example, Akerlof, Dickens, and Perry1996), the graph shows a large spike at zero.

At the employer level, there are two large data sources: the compensation data collected by the Bureau of Labor Statistics (BLS) for the Employment Cost Index, the National Compensation Survey (NCS), and the data collected by the BLS for the Occupational Employment Statistics (OES) survey. Several researchers have used establishment data to examine wage rigidities. Each of these measures has its drawbacks and strengths. A notable example is the study by Lebow, Saks, and Wilson (2003), which finds evidence of significant downward nominal wage rigidity using NCS data. One advantage of the NCS dataset is that it covers the entire compensation package for a job. The drawback is that the sample is relatively small. In the NCS, the BLS surveys firms for a random selection of four to eight occupations in each establishment and obtains comprehensive information only for those occupations. Information about the wage bill of the firm or the industry is not collected.

The OES, on the other hand, is a more complete survey of wages in these establishments. The survey collects information for all workers employed in each job category. Unfortunately, many details about the compensation package for each worker are omitted. For example, the OES does not collect the exact wage; rather it asks for each worker's the wage interval. Essentially the survey provides a sequence of wage intervals and asks how many workers in a given job are in each interval.[38] Not only does it fail to get the exact wage, but because of the large number of firms and individuals, it surveys establishments over a three-year cycle and uses a weighted average to calculate the wages and employment associated with a given job. As a result, the OES measure should be slow to register wage changes.

Figure 14 presents the distribution of wage changes by job in the OES and the distribution of wage changes for employed workers in the OES. A job is defined as one of approximately 800 occupations in one of 300 industries. The wage provided is for each occupation in a given industrial classification; for example, the wages of "research assistants employed by central banks" might be

[38] Averaging wages at different wage intervals could understate the extent of downward nominal wage rigidity. The direction of the averaging effect is unclear a priori, but it should be kept in mind as Wilson (1999) found less rigidity for job averages than for individuals when examining compensation data for a small number of business establishments.

included. The firm-level wage distributions shown in the top panel of figure 14 provides a stark contrast to the distributions found at the individual level. The OES jobs data for 2008–2009 do not show an obvious asymmetry in the distribution of wage changes, nor a spike at zero. Because the jobs that are experiencing an actual wage decline may have a disproportionately small number of workers, the bottom panel of figure14 shows the employment frequency of each wage change. While the variance of the distribution declines when measured by employment, its basic balance remains about the same. The negative tail is substantial and similar to the positive tail, and there is again no large spike at zero.

This evidence from OES data is suggestive of some flexibility in firms' labor costs. It is by no means inconsistent with the presence of downward nominal wage rigidity at the individual level, but it raises the issue of whether firms are able to affect their wage bill via compositional changes in their labor force, downward nominal wage rigidity notwithstanding. Needless to say, much more work is needed to substantiate such a claim. In particular, while the distributions shown in figure 14 indicate that compositional changes can drive wage changes, it still remains to be proven that firms in a recession can deliberately change the composition of their workers so as to lower unit labor costs. If this were indeed the case, it would also be important to know through which channels firms can manage the composition of their workforce, since for example a practice of laying-off more expensive senior workers to hire younger workers at a lower pay would be against the law. Still, more attention should be devoted to analyzing the firm's wage bill rather than individual wages, as compensation for jobs is the relevant unit of labor input when a firm evaluates its costs. In this respect, the information contained in the OES data could prove valuable to examine the effect of job changers on wage and price dynamics.

7. Conclusions

The recent decline of inflation in the United States raises a number of questions of considerable relevance to policymakers. Will inflation continue to fall, or rise towards the central bank's implicit inflation goal? Will well-anchored inflation expectations mitigate or completely offset other disinflationary forces? If U.S. inflation falls, will it behave as if it has a lower bound, as has been the case for Japan? If so, why? To what extent is the experience of Japan a useful guide for current the U.S. experience? What models best serve to capture the dynamics of U.S. inflation in current circumstances of low inflation?

The paper provides partial answers to these questions. In sum, the paper first suggests that long-run inflation expectations, while reasonably well-anchored in both the United States and Japan, do not appear to exert a direct influence on the evolution of Japanese core inflation over the past 20 years. It may be the case that anchored long-run expectations have prevented inflation from spiraling downward in Japan, but we have found it difficult to develop empirical evidence in favor of that proposition.

Second, one-year survey expectations do appear to act as a significant determinant of inflation in both the United States and Japan. In both countries, those expectations respond gradually to the estimated output gap and to recent inflation. In the United States, short-run inflation expectations appear to be anchored in a meaningful sense by long-run expectations, which recently have centered on 2 percent—a reasonable estimate of the Fed's inflation goal. Taken together, these influences can lead to a very slow adjustment of inflation to output and marginal cost: as the output gap widens, one-year expectations fall, pulling down actual inflation. Even as the output gap improves, the influence of recent inflation on expectations continues to depress expectations, in turn slowing inflation on its eventual return towards the central bank's inflation goal. In the Japanese experience, and in the simulation for the United States, this adjustment can take a decade or more.

Because the process that generates one-year expectations is a reduced form, the models estimated on Japanese and U.S. data do not have the status of more theoretically grounded, micro-founded models of inflation. But the models appear to fit the historical data quite well, appear reasonably stable (given the limitations to testing stability inherent in our relatively short samples), and for the samples examined empirically dominate the purely forward-looking models, the hybrid rational expectations models, the more optimistic nonlinear Phillips curve of section 2, and the old-style accelerationist Phillips curves. Thus, the implications of the models based on recent Japanese and U.S. data should be taken as serious cautionary tales for current U.S. policy.

With regard to the lower bound on inflation, the paper briefly examines evidence regarding the extent to which downward nominal wage rigidity—with respect to a firm's wage bill, rather than to individual wages—might slow or stop the decline of inflation. While our data are not completely up to the task, the evidence so far suggests no obvious downward rigidity in the wage bill of the firm, despite widespread evidence indicating downward rigidity in individual wages. Thus one may not be able to take too much comfort in the buffer to disinflation provided by this type of downward nominal wage rigidity.

Overall, we interpret this evidence as a cautionary tale for policymakers. While one should not conclude that inflation must fall, or that deflation is inevitable, the evidence that anchored expectations or wage rigidities will halt the decline of inflation is also not entirely compelling. Thus the risk of further declines in U.S. inflation is a risk that merits serious attention.

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Data Appendix for Sections 3, 4 and 5

Japanese Data		
Variable	**Definition and Source**	**Mnemonic**
Core inflation	Consumer price index less food and energy, 400 times the log change in the price index, Min. of Intern. Affairs & Communic.	π_t
Total inflation	Consumer price index, 400 times the log change in the price index, Min. of Intern. Affairs & Communic.	π_t^T
Output gap	As computed by the Japanese Cabinet Office	$\tilde{y}_t$
Real marginal costs	Labor share of income (equivalently, nominal unit labor costs divided by the price level), OECD	s_t
Exchange rate	Broad nominal exchange rate, JP Morgan	-
Relative price of imported energy	Petroleum, coal and gas import price index, log difference between index and Japanese total CPI, Bank of Japan	rp_t
1-yr. inflation forecast	Consensus Forecasts, one-year ahead forecast, *t* is date forecast is made, semi-annual frequency	π_t^{1y}
6-10-yr. inflation forecast	Consensus Forecasts, 6-10-year ahead forecast, *t* is date forecast is made, semi-annual frequency	π_t^{LR}
U.S. Data		
Core inflation	Consumer price index less food and energy (BLS), annualized quarter-to-quarter percentage change	π_t
Total inflation	Consumer price index (BLS), annualized quarter-to-quarter percentage change	π_t^T
Output gap	One hundred times the log difference between real GDP (BEA) and the HP-filtered trend in real GDP (λ=1600)	$\tilde{y}_t$
1-yr. inflation forecast	Survey of Professional Forecasters, four-quarter-average inflation forecast (FRB Philadelphia), *t* is date forecast is made	π_t^{1y}
10-yr. inflation forecast	Survey of Professional Forecasters, average over next ten years inflation forecast (FRB Philadelphia), *t* is date forecast is made	π_t^{LR}
Real marginal costs	Labor share of income (BLS), computed as 100 times log nominal unit labor cost less log of the implicit price deflator, nonfarm business.	s_t

Note: All data, with the exception of the Japanese output gap, the SPF and Consensus inflation forecasts, are obtained from Haver Analytics.

Table 1 U.S. Inflation, 2008 Peak through October 2010		
Inflation measure	**Change since 2008 peak (percentage points)**	**Inflation rate (12-mo. or 4-qtr. chg.)**
Core CPI (monthly)	-1.94	0.59
Headline CPI (monthly)	-4.35	1.17
Core PCE (monthly)	-1.70	0.94
Headline PCE (monthly)	-3.26	1.26
GDP deflator (quarterly)	-1.37	1.23
Cleveland Fed trimmed mean (monthly)	-2.84	0.77
Cleveland Fed weighted median (monthly)	-2.70	0.51
ECI private compensation (quarterly)	-1.27	1.91

Note: The table documents recent declines in U.S. inflation according to different inflation measures, from their peak values in 2008 through October 2010.
Source: BLS, BEA, and Federal Reserve Bank of Cleveland.

Table 2
The Japanese Phillips Curve: Test for a Shift in Coefficients

$$\pi_t = \sum_{i=1}^{3} a_i \pi_{t-i} + \sum_{j=1}^{2} b_j (y_{t-j} - y^*_{t-j}) + \sum_{k=0}^{1} c_k rp_{t-k}$$

	1971:Q1-1989:Q4		1971:Q1-2009:Q4	
Variable	**Coefficient**	***p*-value**	**Estimated shift, 1990:Q1 to end of sample**	***p*-value**
Lagged inflation ($\sum a_i$, constrained)	1	-	-	-
Output gap ($\sum b_j$)	0.39	0.00	-0.38	0.00
Rel. price of imported energy goods ($\sum c_k$)	2.19	0.008	-0.18	0.003
p-value for unit sum restriction	0.35		0.0051	
R^2	0.83		0.63	
Standard error of regression	1.85		0.78	

Note: The dependent variable is core inflation. The regression is estimated via ordinary least squares, and standard errors are corrected for heteroskedasticity and autocorrelation. The sample period over which the Phillips curve is estimated is 1971:Q1 to 2009:Q4, with the regression allowing for a shift in the coefficients on the output gap and the relative price of energy goods from 1990:Q1 on. All variables are defined in the Data Appendix.

Table 3
The Japanese Phillips Curve with Survey Expectations

$$\pi_t = a\pi_t^{1y} + b\pi_t^{LR} + c\pi_{t-1} + es_t$$
$$[a + b + c = 1]$$

Coefficient	Estimate	*p*-value
1-yr. expectation (a)	0.77	0.00
LR expectation (b)	0.017	0.92
Lagged inflation (c)	0.21	0.041
Marginal costs (e)	0.16	0.007
R^2 : 0.79 *p*-value for unit sum restriction: 0.14		
Excluding the Long-Run Survey Expectation (*b*=0)		
Coefficient	**Estimate**	***p*-value**
1-yr. expectation (a)	0.79	0.00
Lagged inflation (c)	0.21	0.037
Marginal costs (e)	0.16	0.005
R^2 : 0.79 *p*-value for unit sum restriction: 0.31		

Note: The dependent variable is core inflation. The regression is estimated via ordinary least squares, and standard errors are corrected for heteroskedasticity and autocorrelation. The data frequency is semi-annual, and the sample period is 1990:H1 to 2009:H2. All variables are defined in the Data Appendix.

Table 4
A Hybrid Phillips Curve Model for Japan

$$\pi_t = aE_t\pi_{t+1} + b\pi_t^{1y} + (1-a-b)\pi_{t-1}^T + cs_t + c_0$$

$$\pi_t^{1y} = d\pi_{t-1}^{1y} + e\pi_t + f\pi_{t-1} + g\tilde{y}_t + g_0$$

$$\pi_t^T - \pi_{t-1}^T = j(\pi_{t-1}^T - \pi_{t-1}) + k(\pi_{t-1} - \pi_{t-2})$$

$$s_t = \sum_{i=1}^{2} \mathrm{B}_i X_{t-i}$$

$$\tilde{y}_t = \sum_{i=1}^{2} \Gamma_i X_{t-i}$$

Coefficient	Estimate	Standard Error	*t*-statistic
	Inflation equation		
Rational expectation (a)	0.087	0.18	0.47
Survey expectation (b)	0.70	0.24	2.9
Lagged total inflation (1- a - b)	0.21	0.20	1.0
Marginal costs (c)	0.17	0.076	2.3
Intercept (c_0)	-1.1	0.43	-2.6
	Survey expectation equation		
Lagged survey (d)	0.39	0.19	2.0
Current inflation (e)	0.25	0.23	1.1
Lagged inflation (f)	0.16	0.14	1.1
Output gap (g)	0.032	0.052	0.6
Intercept (g_0)	0.22	0.20	1.1
	Error-correction equation for total and core inflation		
Error-correction coeff. (j)	-0.43	0.39	-1.1
Lagged change in inflation (k)	-0.75	0.72	-1.0
Log-likelihood: 138.25			

Note: The dependent variable is core inflation. The data frequency is semi-annual, and the sample period is 1990:H1 to 2009:H2. The parameters are estimated via full-information maximum likelihood, and standard errors are computed using a Berndt-Hall-Hall-Hausmann (BHHH) algorithm. All variables are defined in the Data Appendix.

Table 5

A. **The U.S. Phillips Curve with Survey Expectations**

$$\pi_t = a\pi_t^{1y} + b\pi_{t-1} + cs_t + c_0$$

$$[a + b = 1]$$

$$\pi_t^{1y} = \sum_{i=0}^{4} d_i \pi_{t-i} + \sum_{j=0}^{2} e_j \tilde{y}_{t-j} + e_0$$

Coefficient	**Estimate**	**Standard error**	***p*-value**
	Core inflation equation (3.2)		
Survey expectation (*a*)	0.70	0.12	0.00
Lagged inflation (*b*)	0.30	0.12	0.013
Marginal costs (*c*)	0.053	0.028	0.071
Intercept (c_0)	-0.22	0.093	0.022
R^2=0.69 Standard error of regression: 0.58			
	One-year expectation equation (3.4)		
Current and lagged inflation ($\sum d_i$)	0.66	0.032 (20.4)*	0.00**
Current and lagged output gap ($\sum e_j$)	0.036	0.032 (1.1)*	0.00**
Intercept (e_0)	0.94	0.087	0.00

R^2=0.84
Standard error of regression: 0.26
* *t*-stat. for sum of coefficients
** *p*-value for joint significance of contemporaneous and lagged values

B. **A Hybrid Phillips Curve Model for the United States**

$$\pi_t = aE_t\pi_{t+1} + b\pi_t^{1y} + (1-a-b)\pi_{t-1} + cs_t + c_0$$

$$\pi_t^{1y} = \sum_{i=0}^{4} d_i \pi_{t-i} + \sum_{j=0}^{2} e_j \tilde{y}_{t-j} + e_0$$

$$s_t = \sum_{i=1}^{2} \mathrm{B}_i X_{t-i}$$

$$\tilde{y}_t = \sum_{i=1}^{2} \Gamma_i X_{t-i}$$

Coefficient	**Estimate**	**Standard error (BHHH)**	***t*-statistic**
	Inflation equation		
Rational expectation (*a*)	0.00	0.018	0.056
Survey expectation (*b*)	0.74	0.13	5.6
Lagged inflation (*1- a - b*)	0.26	0.034	7.7
Marginal costs (*c*)	0.048	0.039	1.2
Intercept (c_0)	-0.19	0.060	-3.2

	One-year expectation equation		
Current and lagged inflation ($\sum d_i$)	0.69	0.047 (14.6)*	0.00**
Current and lagged output gap ($\sum e_j$)	0.056	0.038 (1.5)*	0.00**
Intercept (e_0)	0.89	0.99	0.90

Log-likelihood: 254.94
*t-stat. for sum of coefficients
**p-value for joint significance of contemporaneous and lagged values

Note: The dependent variable is core CPI inflation. The data frequency is quarterly, and the sample period is 1990:Q1 to 2010:Q2. The parameters in the top panel are estimated via ordinary least squares, and standard errors are corrected for heteroskedasticity and autocorrelation. The parameters in the bottom panel are estimated via full-information maximum likelihood, and the standard errors are computed via a Berndt-Hall-Hall-Hausmann (BHHH) algorithm. All variables are defined in the Data Appendix.

Table 6
The Evolution of One-Year Expected U.S. Inflation

$$\pi_t^{1y} = \sum_{i=0}^{4} d_i \pi_{t-i} + f\pi_t^{LR} + \sum_{j=0}^{2} e_i \tilde{y}_{t-i} + e_0$$

$$[\sum d_i + f = 1]$$

Coefficient	Estimate	Standard error	p-value
Current and lagged inflation ($\sum d_i$)	0.19	0.048 (3.9*)	0.00**
Long-run inflation expectation (f)	0.81	0.048	0.00
Current and lagged output gap ($\sum e_j$)	0.13	0.023 (5.5*)	0.00**
Intercept (e_0)	-0.12	0.022	0.00

R^2=0.92
Standard error of regression: 0.19
p-value for unit sum constraint: 0.66
*t-stat. for sum of coefficients
**p-value for joint significance of contemporaneous and lagged values

Note: The dependent variable is the one-year inflation forecast from the Survey of Professional Forecasters. The data frequency is quarterly, and the sample period is 1990:Q1 to 2010:Q2. The parameters are estimated via ordinary least squares, and standard errors are corrected for heteroskedasticity and autocorrelation. All variables are defined in the Data Appendix.

Table 7
Non-nested Hypothesis Tests of U.S. Models from Sections 2 and 4

$$y_{Pt} = \hat{\beta}_P X_{Pt} + \hat{\lambda}_P \hat{F}_{Jt} + \varepsilon_{Pt}$$

$$y_{Jt} = \hat{\beta}_J X_{Jt} + \hat{\lambda}_J \hat{F}_{Pt} + \varepsilon_{Jt}$$

P: Nonlinear Phillips curve, *J*: Japanese-style survey expectations model

End Date	*p*-value of $\hat{\lambda}_P$, null = Nonlinear Phillips Curve	*p*-value of $\hat{\lambda}_J$, null = Japanese-style survey expectations model
2005:Q4	0.000	0.21
2007:Q4	0.000	0.28
2009:Q4	0.000	0.19
2010:Q2	0.000	0.10

Note: The table reports the Davidson-MacKinnon (1981) test for model specification. The test involves the regression of the dependent variable *y* in one model on the regressors *X* of that model, along with the fitted values $\hat{F}$ of the alternative model. The same type of regression is also run for the alternative model. Data are quarterly, from 1991:Q1 to selected end dates.

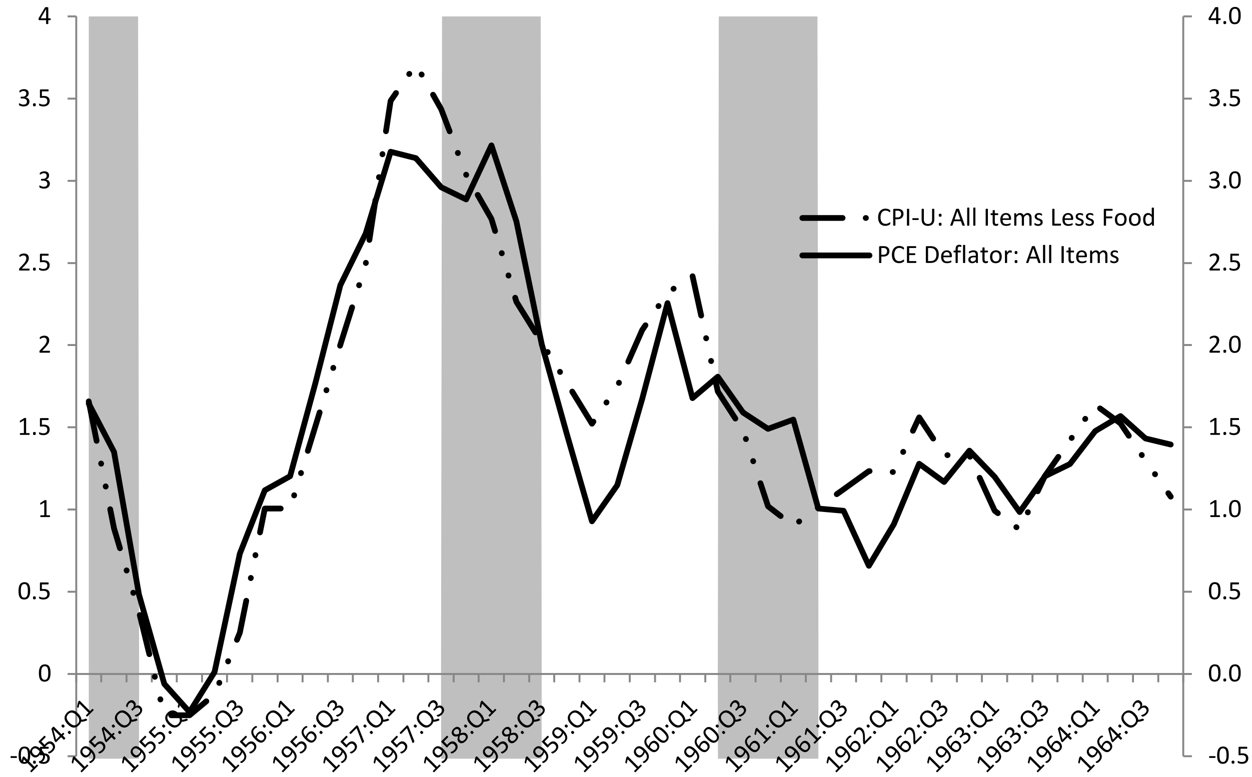


Fig. 1 U.S. Inflation, 1954:Q1 to 1964:Q4

Source: BLS and BEA.

Note: The figure plots four -quarter inflation for different price indices. Inflation at date *t* is defined as $\pi_t^4 = 100 * \left(\frac{P_t}{P_{t-1}} - 1 \right)$, where P_t is the price index in quarter *t*. The shadings represent NBER-defined recession periods.

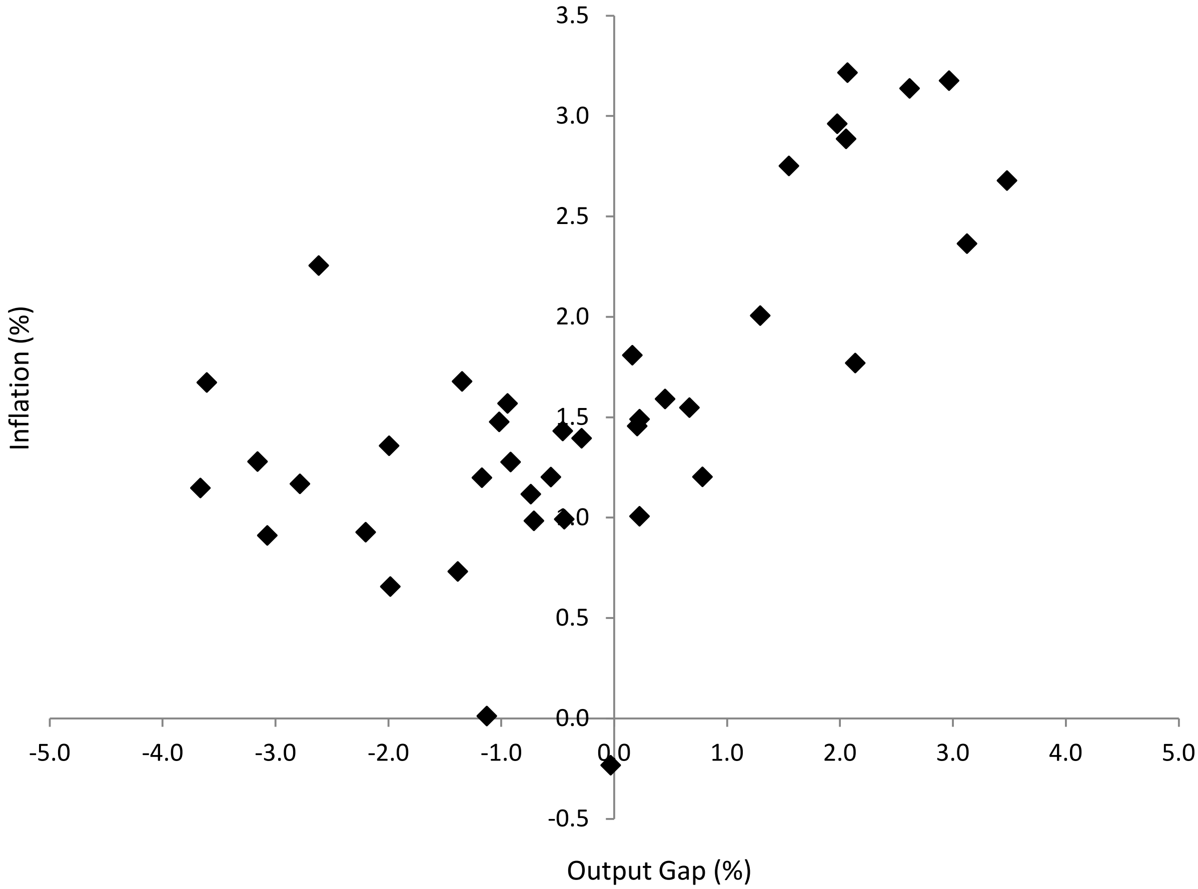


Fig.2 U.S. Output Gap and Inflation, 1954:Q1 to 1963:Q4

Source: BEA, CBO, and authors' calculations.

Note: The figure plots the output gap against four-quarter-ahead inflation, defined as $\pi_{t+4}^{4} = 100 * \left(\frac{P_{t+4}}{P_t} - 1 \right)$, with P_t denoting the PCE (all items) deflator. The output gap estimate is from the Congressional Budget Office, and it is depicted as a four -quarter moving average with declining weights to smooth out some short-term volatility.

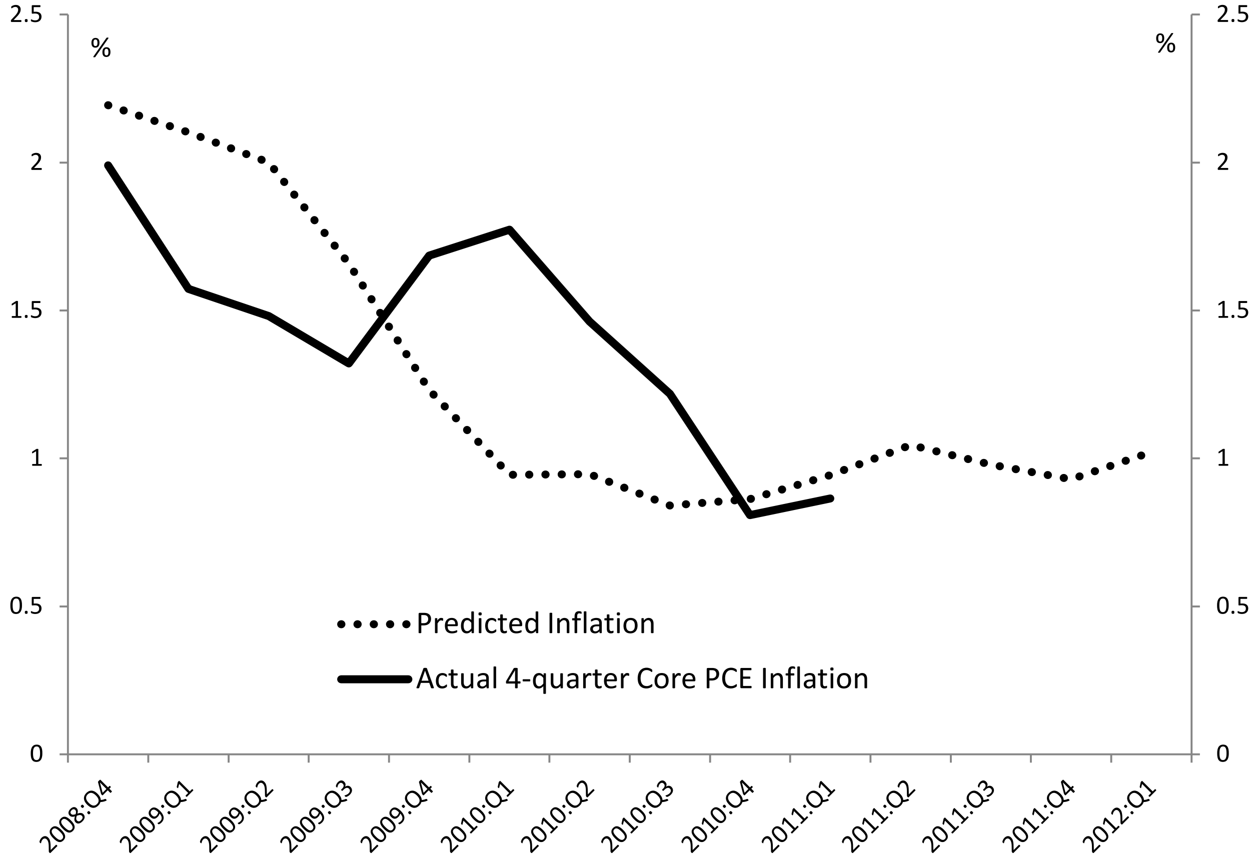


Fig. 3 Actual and Predicted U.S. Inflation, 2008:Q4 to 2012:Q1

Source: BEA and authors' calculations.

Note: The figure shows actual four -quarter inflation as measured by the PCE (all items less food and energy) and its predicted value according to the nonlinear Phillips curve (2.1) in the text, which is estimated over the period 1954:Q1 to 1963:Q4. In the simulation, the output gap takes the value estimated by the Congressional Budget Office, and long-run inflation expectations are measured by the Philadelphia survey's median 10-year inflation expectations.

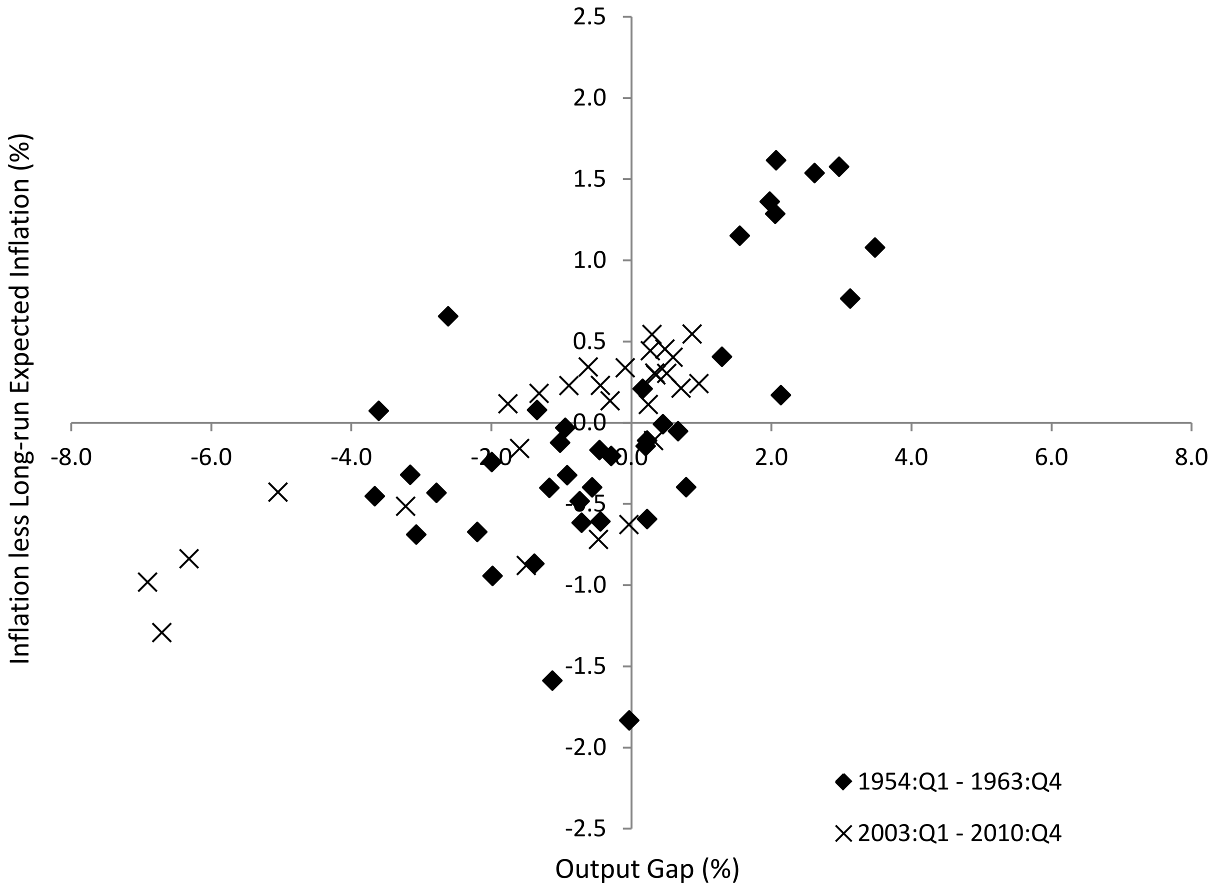


Fig. 4 U.S. Output Gap and Inflation, 1954:Q1 to 1963:Q4 and 2003:Q1 to 2010:Q4

Source: BEA, CBO, Board of the Federal Reserve System, and authors' calculations.

Note: The figure plots the output gap against four -quarter-ahead inflation in deviation from long-run inflation. The deviation of inflation from long-run inflation is given by $\pi_{t+4}^{4} - \pi_t^{LR} = 100 * \left(\frac{P_{t+4}}{P_t} - 1 \right) - \pi_t^{LR}$. Over the period 1954:Q1 to 1963:Q4, the price index P_t is the PCE (all items) deflator and π_t^{LR} is proxied by a constant equal to the average rate of inflation prevailing over this period (1.6 percent). Over the period 2003:Q1 to 2010:Q4, inflation is measured by the core PCE (all items less energy and food) index, and π_t^{LR} is given by the Philadelphia survey's median 10-year inflation expectations.

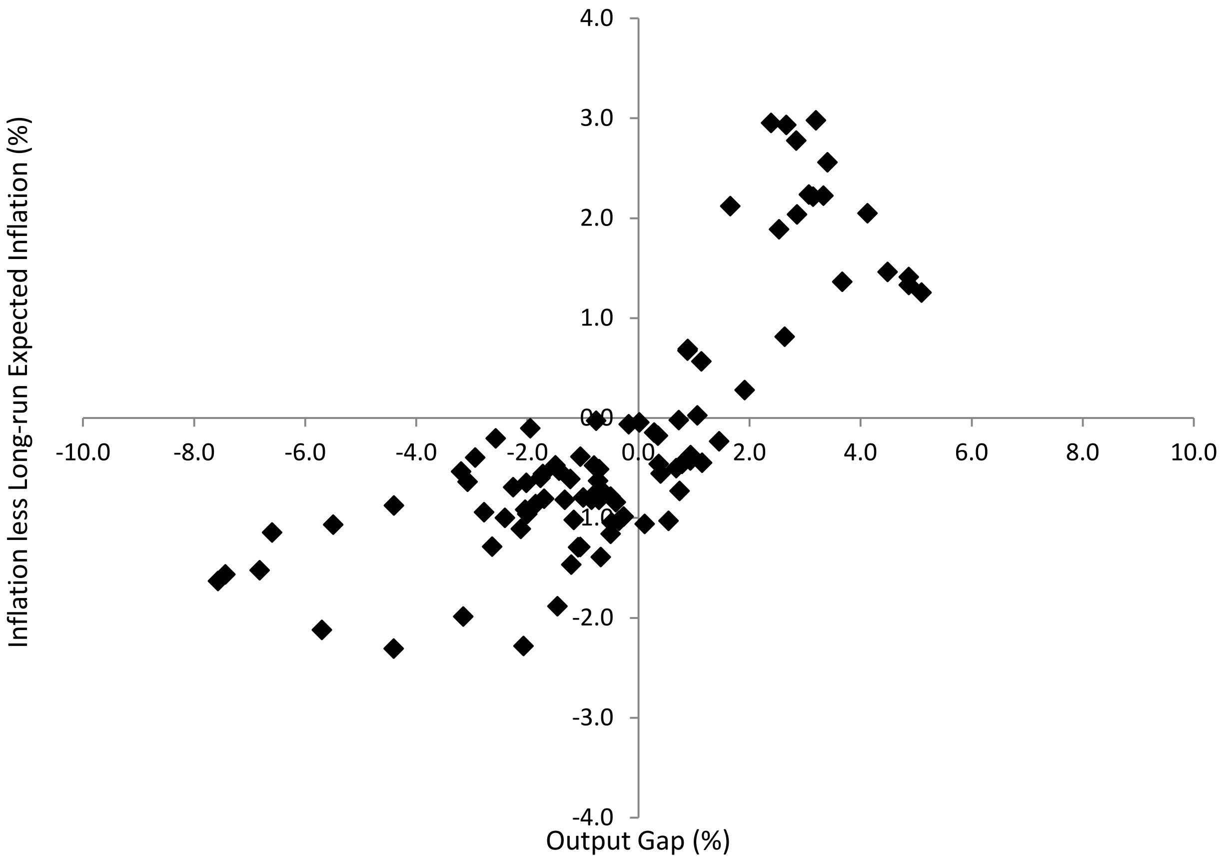


Fig. 5 U.S. Output Gap and Inflation, 1964:Q1 to 1969:Q4 and 1981:Q1 to 1997:Q4

Source: BEA, CBO, Board of the Federal Reserve System, and authors' calculations.

Note: The figure plots the output gap against four -quarter-ahead inflation in deviation from long-run inflation. The deviation of inflation from long-run inflation is given by $\pi_{t+4}^{4} - \pi_{t}^{LR} = 100 * \left(\frac{P_{t+4}}{P_t} - 1 \right) - \pi_{t}^{LR}$. The price index P_t is the core PCE (all items less food and energy) deflator, and π_{t}^{LR} is given by the Philadelphia survey's median 10-year inflation expectations. For the earlier period, this series is not available and it is constructed by fitting the evolution of the 10-year Philadelphia survey measure over the period 1981 to 2006 to that earlier sample.

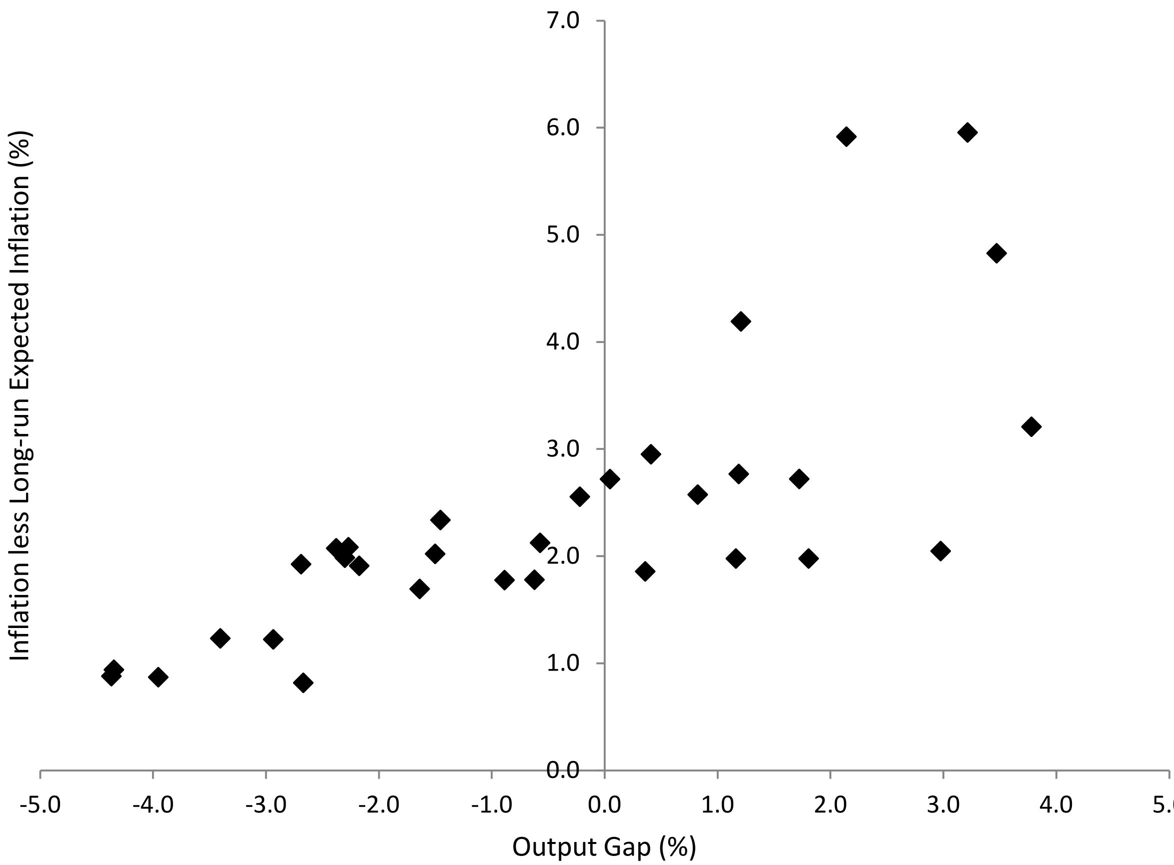


Fig. 6 U.S. Output Gap and Inflation, 1973:Q1 to 1980:Q4

Source: BEA, CBO, Board of the Federal Reserve System, and authors' calculations.

Note: The figure plots the output gap against four -quarter-ahead inflation in deviation from long-run inflation. The deviation of inflation from long-run inflation is given by $\pi_{t+4}^{4} - \pi_{t}^{LR} = 100 * \left(\frac{P_{t+4}}{P_t} - 1 \right) - \pi_{t}^{LR}$. The price index P_t is the core PCE (all items less food and energy) deflator, and π_{t}^{LR} is given by the Philadelphia survey's median 10-year inflation expectations.

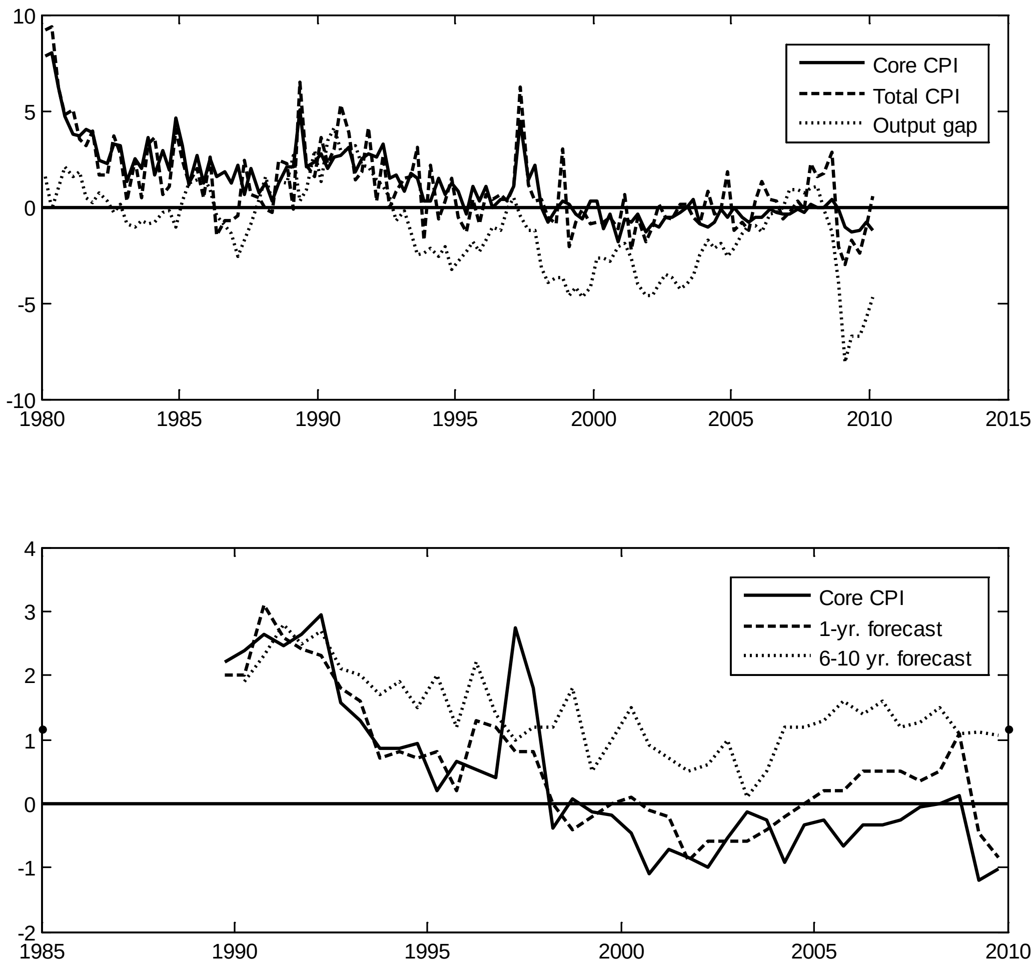


Fig. 7 Key Japanese Data

Note: The first chart in the figure plots quarterly values for inflation and the output gap. The second chart shows semi-annual data for core inflation and Consensus inflation forecasts at different horizons. All variables are defined in the Data Appendix.

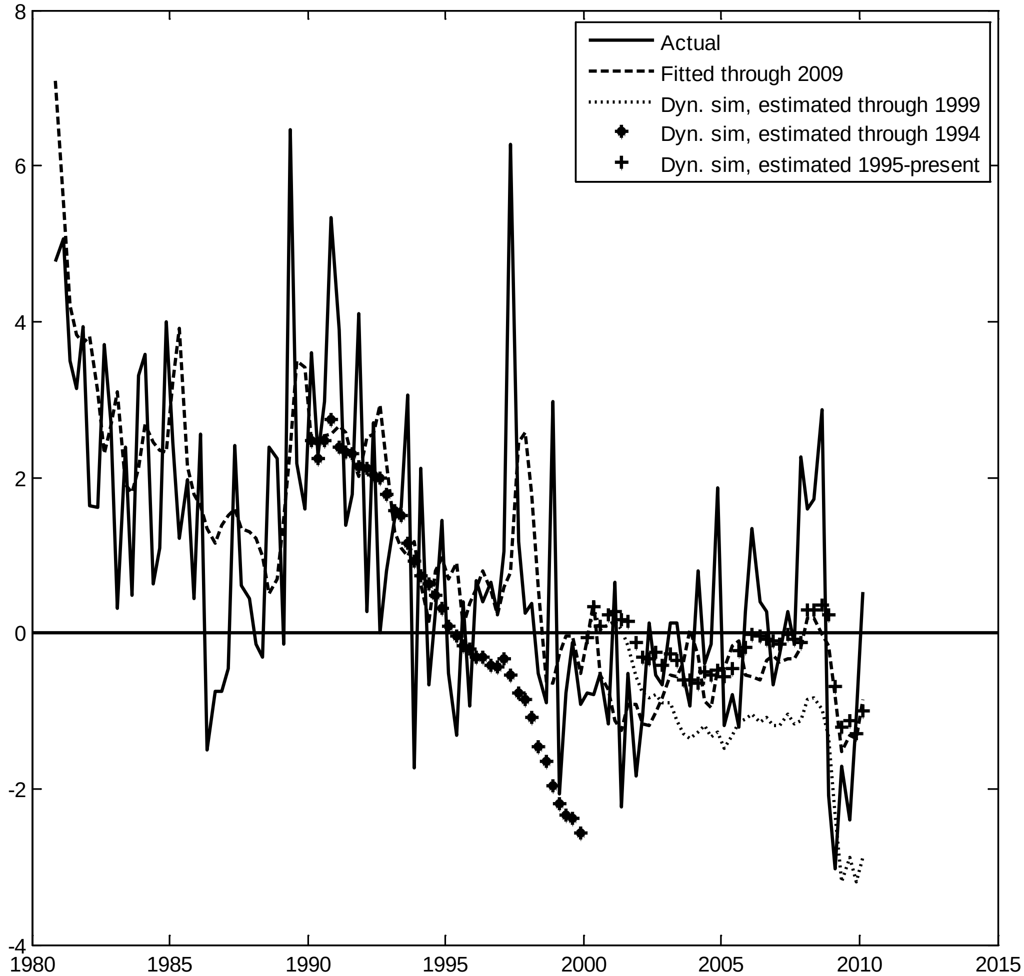
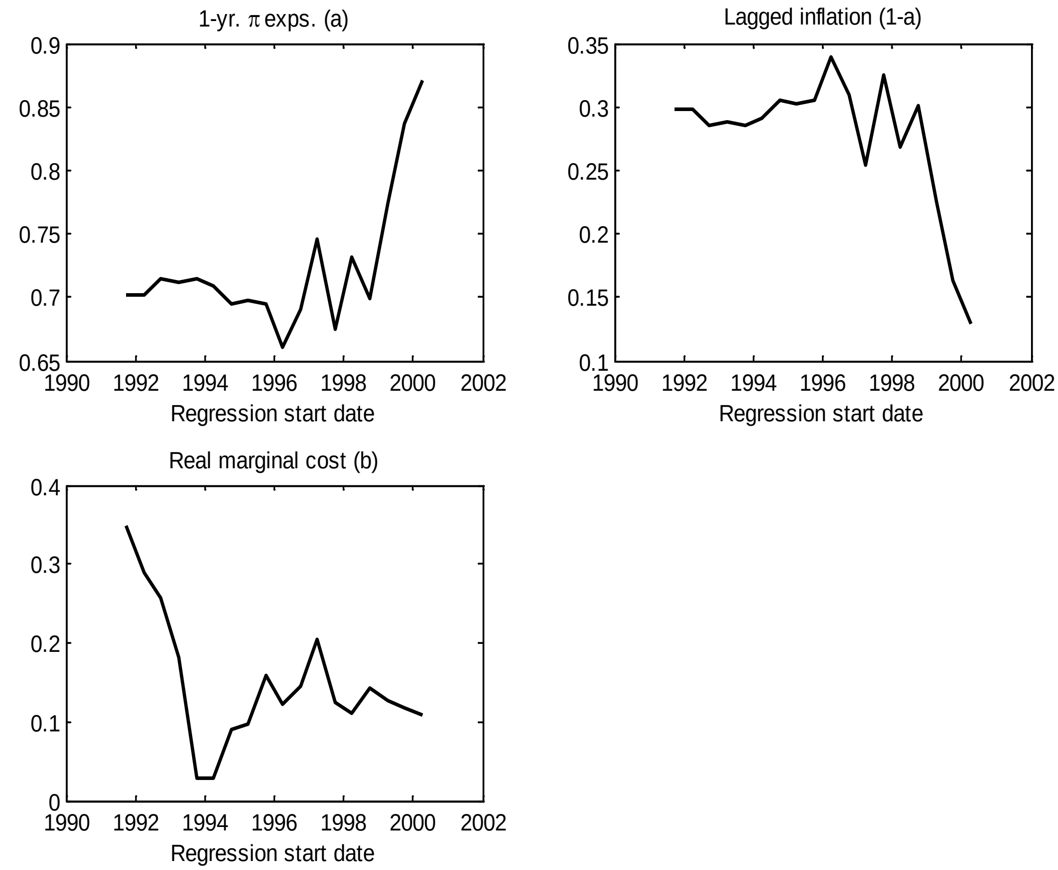


Fig. 8 Fit of Conventional Phillips Curve for Japan, 1980-2010

Note: The figure reports actual core CPI inflation and the predicted values of inflation from a conventional Phillips curve of the form $\pi_t = \sum_{i=1}^{3} a_i \pi_{t-i} + \sum_{j=1}^{2} b_j (y_{t-j} - y^*_{t-j}) + \sum_{k=0}^{1} c_i rp_{t-k}$. The equation is estimated via ordinary least squares for a variety of subsamples as indicated in the legend. In the dynamic simulations, actual values are used for the output gap and the relative price of imported oil. All variables are defined in the Data Appendix.



$$\pi_t = a\pi_t^{1y} + (1-a)\pi_{t-1} + bs_{t-1}$$

Fig. 9 Rolling estimates of Japanese Phillips Curve with One-year Survey Expectations

Note: The charts report coefficient estimates from 10-year-window rolling regressions for the Phillips curve with one-year inflation survey expectations as indicated in the caption. The coefficients are estimated via ordinary least squares. The data frequency is semi-annual and the sample period is 1991:H1 to 2009:H2. The dependent variable is core CPI inflation. All variables are defined in the Data Appendix.

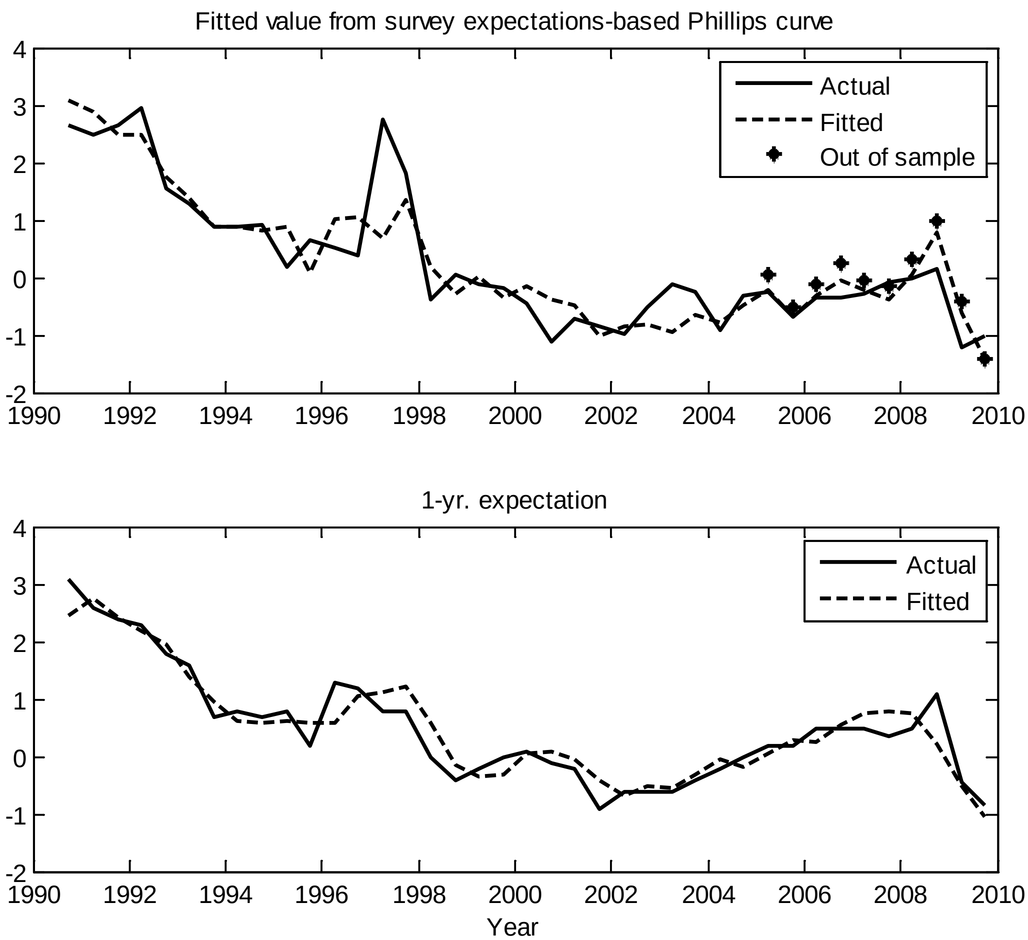


Fig. 10 Japanese Phillips Curve with One-year Survey Expectations

Note: The first chart reports actual and predicted core CPI inflation. In- and out-of-sample predictions are computed from an estimated Phillips curve of the form: $\pi_t = a\pi_t^{1y} + (1-a)\pi_{t-1} + bs_{t-1}$. The second chart shows actual and fitted values for one-year expected inflation, where the fitted equation takes the form: $\pi_t^{1y} = a\pi_{t-1}^{1y} + b\pi_{t-1} + cs_t + d\tilde{y}_t + c_0$. For both equations, the parameters are estimated via ordinary least squares. The data frequency is semi-annual, and the estimation period is 1990:H1 to 2009:H2 for the in-sample forecasting exercise. For the out-of-sample exercise, the estimation period is 1990:H1 to 2004:H2. All variables are defined in the Data Appendix.

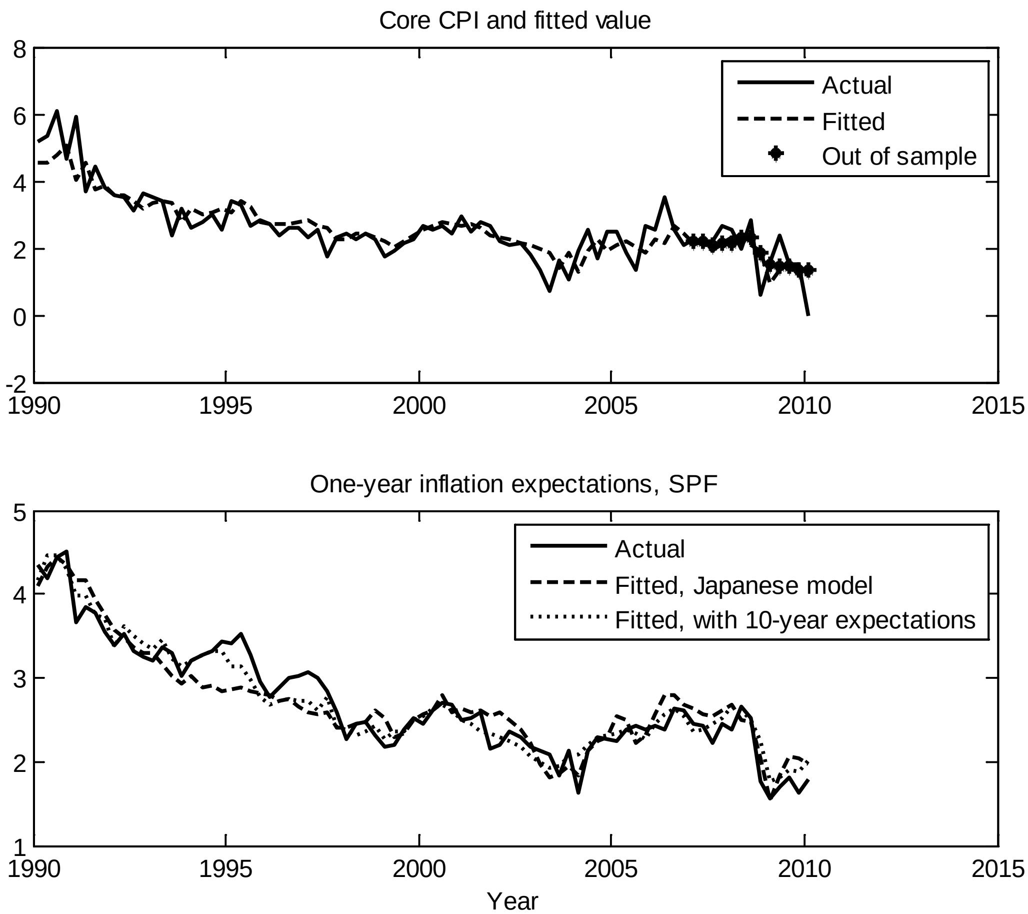


Fig. 11 U.S. Inflation, Actual and Fitted from Japanese Inflation Model

Note: The first chart reports actual and predicted core CPI inflation with the same Phillips curve specification as the one used on Japanese data: $\pi_t = a\pi_t^{1y} + (1-a)\pi_{t-1} + bs_{t-1}$. The second chart shows actual and fitted values for one-year expected inflation, where the fitted equation takes the form: $\pi_t^{1y} = a\pi_{t-1}^{1y} + b\pi_{t-1} + cs_t + d\tilde{y}_t + c_0$. We also consider the fit of an augmented specification that includes as an additional regressor the 10-year Consensus inflation expectations. The coefficients are estimated via ordinary least squares. The data frequency is quarterly, and the estimation period is 1990:Q1 to 2010:Q2 for the in-sample forecasting exercise. For the out-of-sample exercise, the estimation period is 1990:Q1 to 2006:Q4. All variables are defined in the Data Appendix.

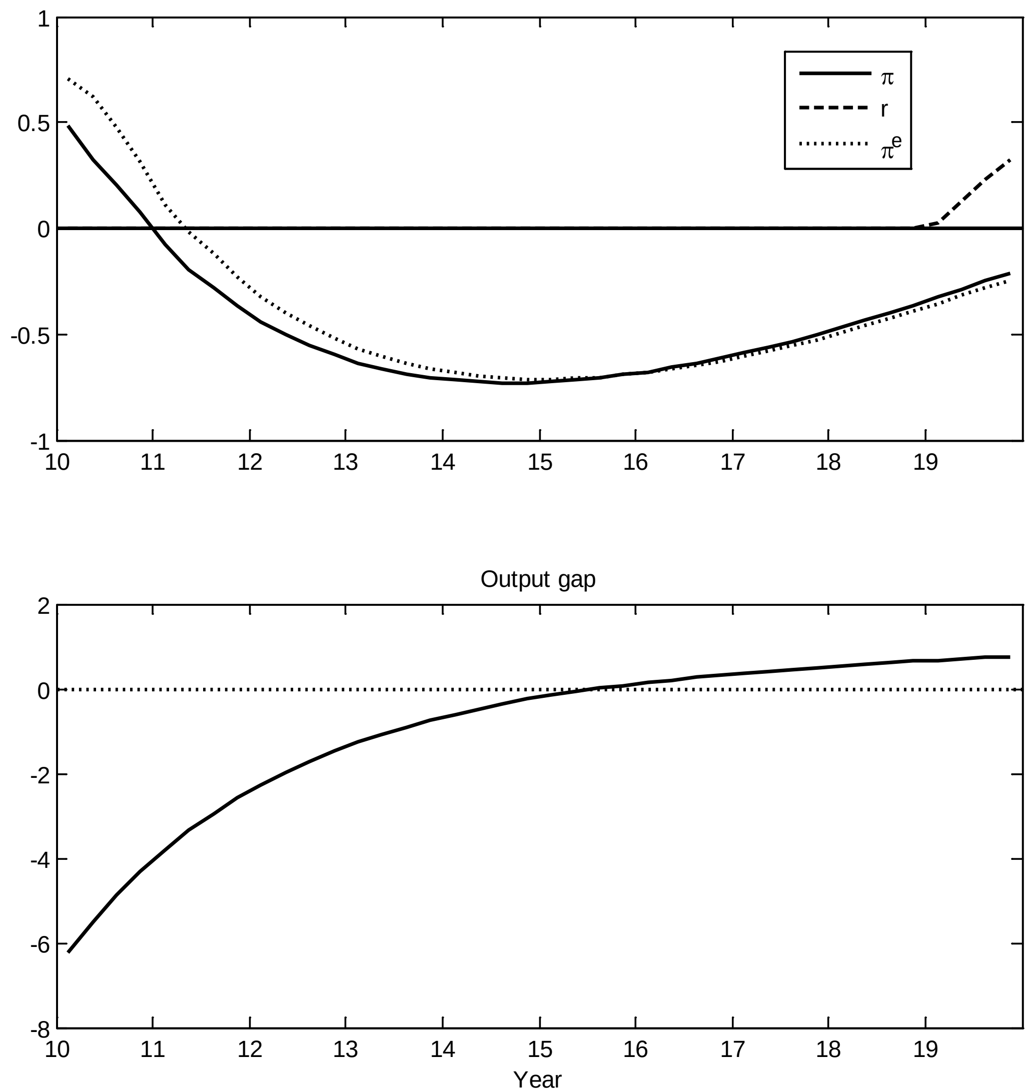


Fig. 12 Simulated Values for U.S. Core Inflation, 2011 to 2020

Note: The charts report simulated values for core inflation, the Federal funds rate, and the output gap using the inflation specification $\pi_t = a\pi_t^{1y} + b\pi_{t-1} + cs_t + c_0$, and $\pi_t^{1y} = \sum_{i=0}^{4} d_i \pi_{t-i} + f\,\pi_t^{LR} + \sum_{j=0}^{2} e_i \tilde{y}_{t-i} + e_0$ as the law of motion for one-year inflation expectations See tables 5 and 6 for estimated parameter values. The model is augmented by a law of motion for the Federal funds rate and the output gap (equations 5.1 and 5.2 in the text), and is initialized with economic conditions prevailing in 2010 as detailed in the text.

Fig. 13 Wage Growth for Job Stayers Over the Period 2005-2007 from PSID Survey

Source: University of Michigan PSID and authors' calculations

Note: The chart reports the distribution of wage changes for people who worked at the same job over the period 2005 to 2007. The data source is the PSID. The chart shows a large spike at zero, consistent with the notion of a significant role for downward nominal wage rigidity.

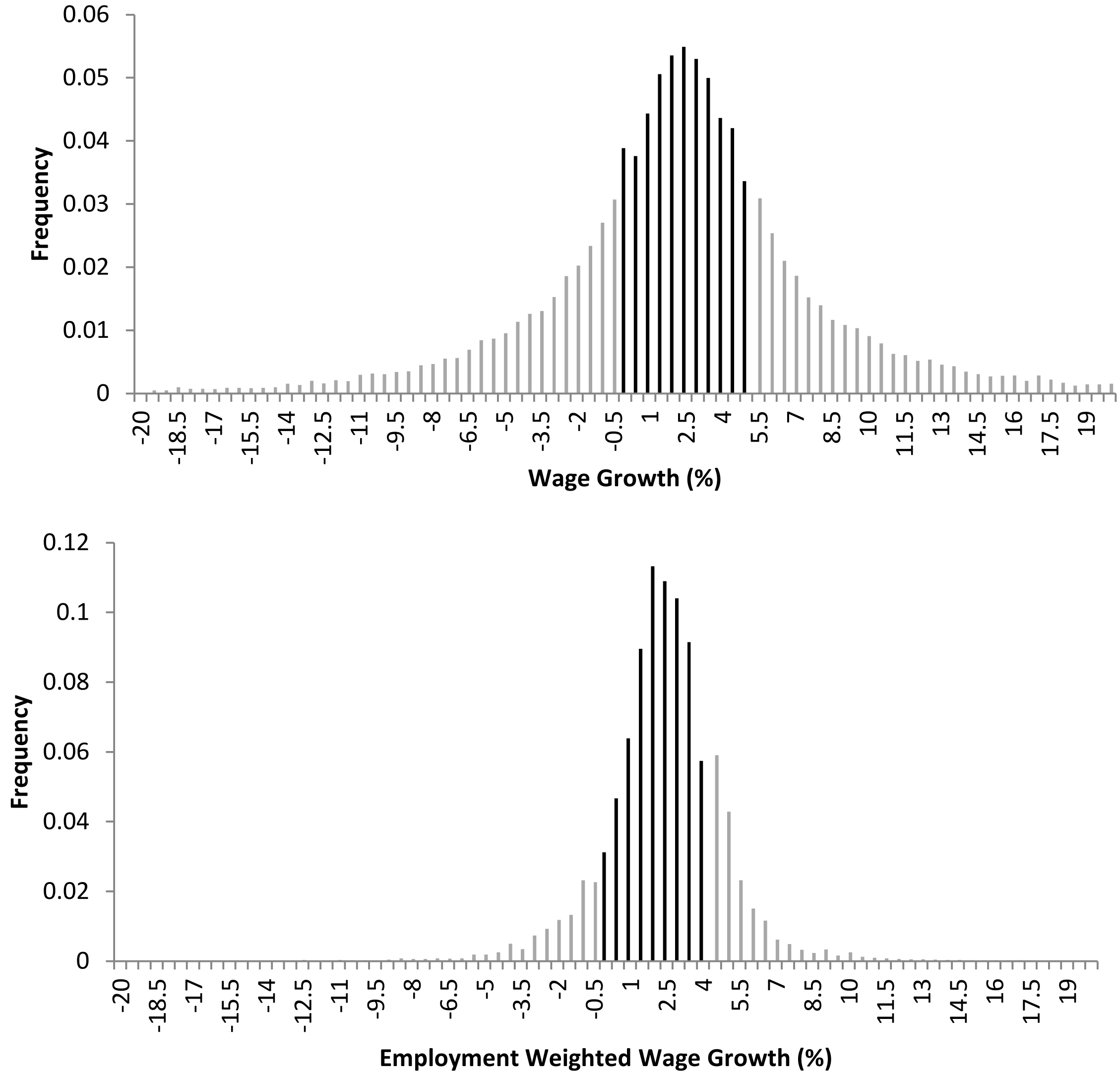


Fig. 14 Wage Growth over the period 2008-2009, Employer Data from OES Survey

Source: BLS and authors' calculations

Note: The charts report the distribution of wage changes over the period 2008-2009. The data source is the OES. The first chart shows the distribution of wage changes for all occupations and industries. The OES categorizes approximately 800 occupations and 300 industries. The second chart weights the wage changes for all occupations and industries by the number of employees in each job category and industry. The darker columns in the charts represent the frequencies of wage changes from zero to 2 times the mode of the distribution. For example, the mode of the distribution in the lower chart is 2, and the darker columns show the frequencies of wage changes over the interval [0,4].

Appendix

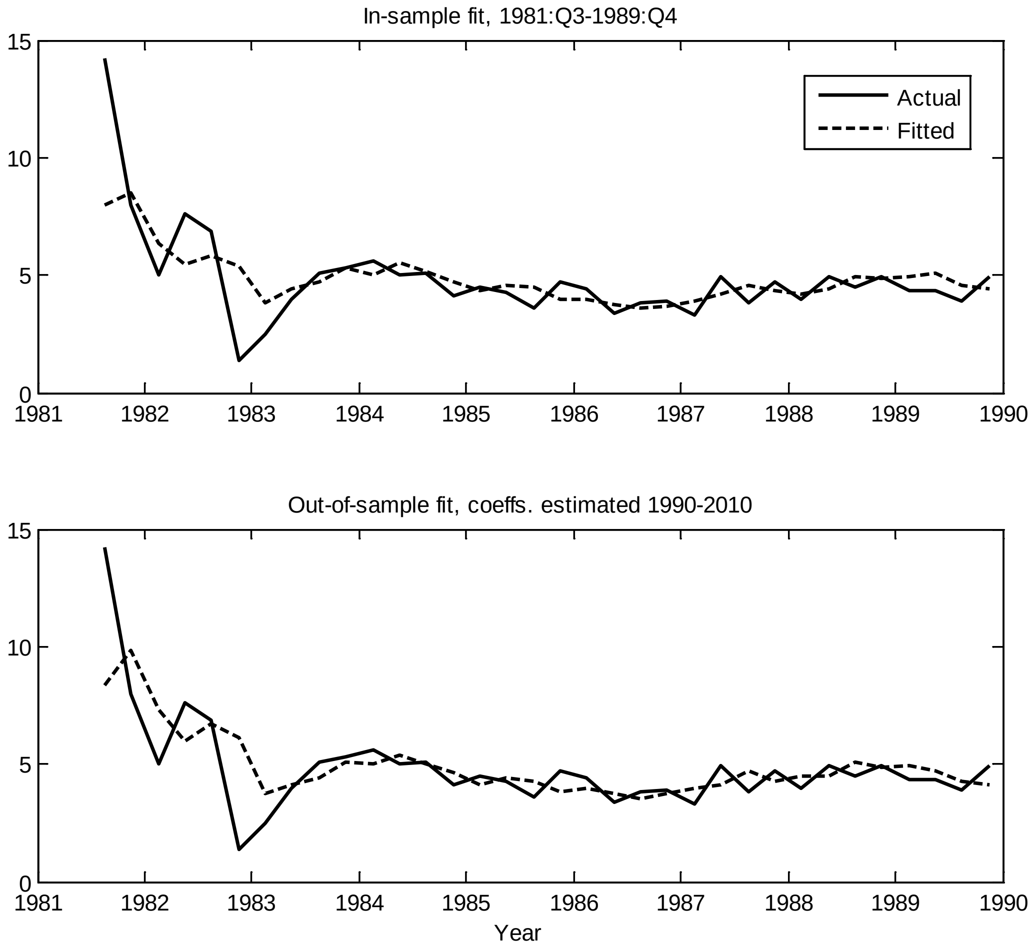


Fig. A1 Fit of Japanese Model of Inflation on pre-1990 U.S. Inflation Data

Note: The first chart shows the in-sample fit from estimating the Japanese model of inflation (see panel *A*. in Table 5) on U.S. data over the period 1981:Q3 to 1989:Q4. The second chart uses the estimates reported in Table 5 over the period 1990:Q1 to 2010:Q2 to assess the out-of-sample fit of the model. Inflation is measured by the core Consumer Price Index. All variables are defined in the Data Appendix.