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U.S. Inflation Dynamics: What Drives Them Over Different Frequencies?

Ravi Balakrishnan and Sam Ouliaris

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U.S. Inflation Dynamics: What Drives Them Over Different Frequencies?

Prepared by Ravi Balakrishnan and Sam Ouliaris[1]

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Abstract

This paper aims to improve the understanding of U.S. inflation dynamics by separating out structural from cyclical effects using frequency domain techniques. Most empirical studies of inflation dynamics do not distinguish between secular and cyclical movements, and we show that such a distinction is critical. In particular, we study traditional Phillips curve (TPC) and new Keynesian Phillips curve (NKPC) models of inflation, and conclude that the long-run secular decline in inflation cannot be explained in terms of changes in external trade and global factor markets. These variables tend to impact inflation primarily over the business cycle. We infer that the secular decline in inflation may well reflect improved monetary policy credibility and, thus, maintaining low inflation in the long run is closely linked to anchored inflation expectations.

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Author(s) E-Mail Address: rbalakrishnan@imf.org; souliaris@imf.org

[1] This paper has benefited from comments by our colleagues in North American Division and Hasan Bakhshi. Errors and omissions are the authors' sole responsibility.

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I. Introduction And Summary

In the United States, as in many other industrialized countries, inflation has remained relatively low since the mid-1990s. This favorable outcome has occurred despite long periods of robust economic growth, accommodative monetary policy, and, more recently, substantial oil and other commodity price increases.

Traditional, backward-looking empirical models for inflation tend to overpredict inflation, particularly after the mid-1990s. Further empirical investigation therefore remains warranted, notwithstanding a broad consensus that improved monetary policy credibility and increased globalization are the most likely reasons behind the decline in inflationary pressures. This paper analyzes the factors driving U.S. inflation by separating structural from cyclical effects and examining the impact of globalization on these separate components of inflation. Using data for 1960–2005, frequency domain techniques identify a smooth, secular decline in inflation that commenced around 1980. They also suggest that there has been a reduction in the size and volatility of the business-cycle component of inflation—a moderation that appears related to the growing impact of globalization.

Two distinct models are used to analyze inflation in the time and frequency domain. The first is the traditional Phillips curve (TPC) model that is largely backward-looking, emphasizing the role of lagged inflation and the output gap. The other model is the "forward-looking" New Keynesian Phillips Curve (NKPC), which places greater emphasis on expected inflation and changes in marginal cost.

Overall, we conclude that the long-run secular decline in inflation cannot be explained in terms of changes in external trade and global factor markets, since these variables tend to impact inflation primarily over the business cycle. We infer that the secular decline in inflation in the United States reflects improved monetary policy credibility and the adoption of an low implicit target for inflation and, thus, maintaining low inflation in the long run is closely linked to anchored inflation expectations.

The organization of this paper is as follows: Section II contains the results (time domain and frequency domain) for the traditional Phillips curve model; the results for the new Keynesian Phillips curve are presented in Section III; and Section IV concludes.

II. Traditional Phillips Curve

A. Time Domain Analysis

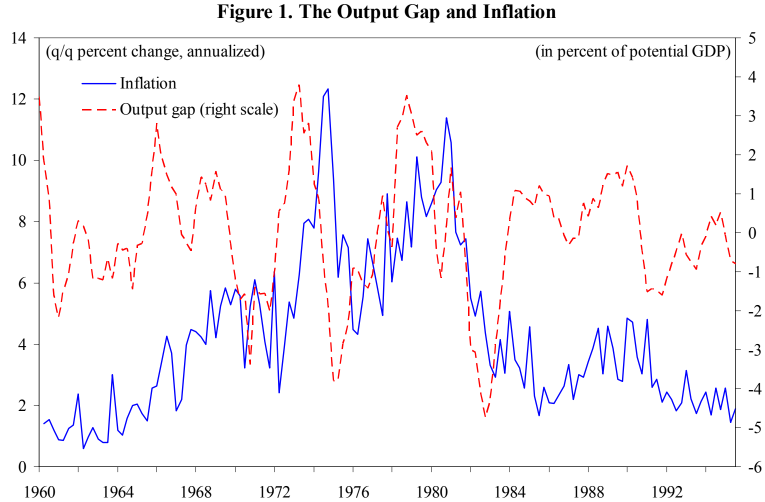
The traditional Phillips curve (TPC) models inflation in terms of excess demand (i.e., the output gap) and lagged values of inflation. Let π_t denote inflation and $\hat{y}_t - y_t^*$ the log deviation of real GDP from its long run trend. A common specification for the traditional Phillips curve is:

$$\pi_t = \sum_{i=1}^{h} \varphi_i \pi_{t-i} + \delta(\hat{y}_{t-1} - y^*_{t-1}) + \varepsilon_t \qquad (1)$$

where ε_t is a random disturbance and h is the number of lagged inflation terms in the model. The restriction $\sum_{i=1}^{h} \varphi_i = 1$ is often imposed so that the model implies no long-run trade-off between output and inflation. Alternative specifications may use different measures of excess demand (e.g., the unemployment rate and capacity utilization).

The bivariate relationship between inflation and the output gap appears to have weakened since 1990. The relationship between the quarterly GDP deflator inflation and the output gap, as measured by a standard Hodrick-Prescott filter, is shown in Figure 1.[2] Prior to 1990, movements in the output gap clearly lead inflation. The empirical relationship between these two variables appears to have weakened since then, however—the statistical correlation between inflation and the output gap is slightly positive (and significant) for 1960–2005, but is negative and insignificant for 1991–2005.

Figure 1. The Output Gap and Inflation



The TPC model provides a reasonable explanation of inflation over 1960–2005. Equation (2) shows the results of estimating (1) using ordinary least squares and four lags of inflation. The regression results, which are consistent with previous estimates reported in the literature (e.g., Rudebusch and Svensson, 1999), imply that a positive (and statistically significant) relationship exists between inflation, past inflation, and the lagged output gap—with the

[2] We focus on GDP deflator inflation. Similar results hold if we use CPI (headline or core) inflation to measure inflation.

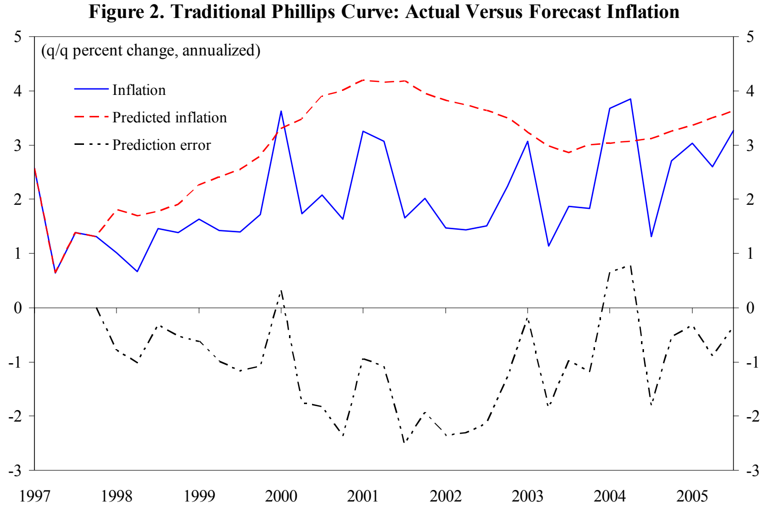
model accounting for over 80 percent of the variation in inflation.[3] Moreover, the sum of the coefficients on the lagged inflation variables is not—in a statistical sense—significantly different from unity, implying the absence of a trade-off between inflation and the output gap in the long run.

$$\pi_t = \underset{(1.9140)}{0.0707} + \underset{(4.1332)}{0.4765}\pi_{t-1} + \underset{(1.0338)}{0.0946}\pi_{t-2} + \underset{(1.4016)}{0.1131}\pi_{t-3} + \underset{(2.9882)}{0.2480}\pi_{t-4} + \underset{(4.4829)}{0.0653}\,(y_{t-1} - y^*_{t-1})$$
$$\bar{R}^2 = 0.8165 \quad Durbin\text{-}Watson = 1.9755 \quad F_{(1,172)}(\sum_1^4 \varphi_i = 1) = 1.8445 \quad (p\text{-}value:\ 0.1762) \qquad (2)$$

The out-of-sample forecasting performance of the TPC model is, however, poor. When equation (2) is estimated over 1960–97 and then used to forecast inflation after 1997, systematic overprediction is evident (Figure 2). This overprediction occurs despite the fact that the estimated coefficients for the subperiod (equation (3)) are similar to the full-period estimates (i.e., there is no evidence of structural change between the two periods). This suggests that the model is accounting for the secular decline only through the lagged inflation terms and, hence, with a one-year (i.e., a four-quarter) delay.[4]

$$\pi_t = \underset{(1,8669)}{0.00783} + \underset{(3.8318)}{0.4883}\pi_{t-1} + \underset{(0.9426)}{0.0981}\pi_{t-2} + \underset{(1.0990)}{0.0998}\pi_{t-3} + \underset{(2.6444)}{0.2428}\pi_{t-4} + \underset{(4.2818)}{0.0677}\,(y_{t-1} - y^*_{t-1})$$
$$\bar{R}^2 = 0.8127 \quad Durbin\text{-}Watson = 1.9676 \quad F(break,\ 1998:1) = 0.3831 \quad (p\text{-}value: 0.8891) \qquad (3)$$

Figure 2. Traditional Phillips Curve: Actual Versus Forecast Inflation



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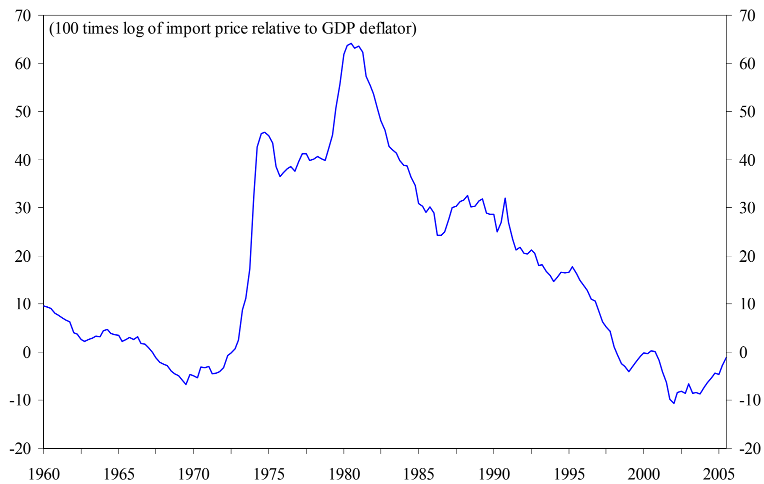
[3] Numbers in parenthesis refer to t-statistics. The sample begins in 1961:Q2 because of the inclusion of 4 lagged inflation terms.

[4] Interestingly, TPC models have the opposite problem for the euro area since the late 1990s—they underpredict inflation (see Bakhshi, 2006).

Incorporating external shocks fails to resolve the overprediction problem. Some researchers have attributed the overprediction problem to the omission of external price shocks that have kept both imported and domestically produced goods inflation comparatively low. Indeed, looking at the trend in the import price deflator in recent years (Figure 3), it is certainly possible for import prices to have put downward pressure on overall inflation, at least until the recent increases in commodity and oil prices occurred.

Figure 3. Import Price Deflator



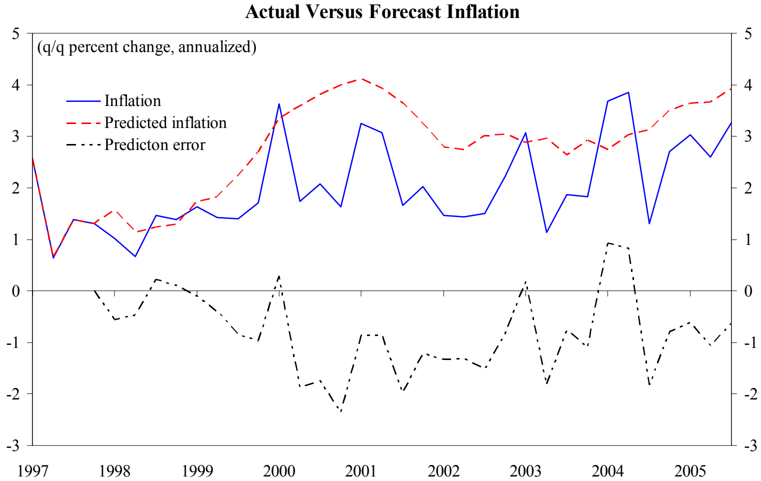
We allow for external shocks by including the terms of trade in the base specification. Doing so (equation (4)) produces a significant positive association between inflation and movements in the terms of trade. While it reduces the degree of overprediction, it still fails to resolve the underlying problem (Figure 4).

$$\pi_t = \underset{(3.9168)}{0.1155} + \underset{(4.3345)}{0.3748}\pi_{t-1} + \underset{(1.0439)}{0.0880}\pi_{t-2} + \underset{(1.8461)}{0.1324}\pi_{t-3} + \underset{(3.5059)}{0.2680}\pi_{t-4} + \underset{(3.6110)}{0.0515}(y_{t-1} - y^*_{t-1})$$

$$+ \underset{(3.2411)}{0.04560}\Delta tot_{t-1} + \underset{(2.0282)}{0.0320}\Delta tot_{t-2} + \underset{(2.3064)}{0.0319}\Delta tot_{t-3} + \underset{(0.2490)}{0.0043}\Delta tot_{t-4} \quad (4)$$

$$\overline{R}^2 = 0.8449 \quad Durbin\text{-}Watson = 1.9922 \quad F(break,\ 1998:1) = 0.4572 \quad (p-value: 0.9151)$$

Figure 4. Traditional Phillips Curve with Terms of Trade Shocks: Actual Versus Forecast Inflation



B. What Causes the Overprediction in TPC Models?

The overprediction problem strongly suggests that the TPC model is misspecified. To clarify potential sources of overprediction, consider the following version of the expected-augmented Phillips curve:

$$\pi_t = \pi_t^e + \lambda(y_{t-1} - y_{t-1}^*) + s_t + \varepsilon_t \tag{5}$$

where π_t^e is the inflation expectation at time t, and s_t is a supply shock. The standard TPC model with terms of trade shocks can be thought of as using a weighted average of lagged inflation to proxy inflation expectations. Apart from the fact that this is purely backward looking, it is clear from the Lucas critique that the way in which these expectations are formed could change as the policy regime shifts. Many analysts have argued that such changes have indeed taken place in the United States, with monetary policy credibility having increased considerably since the early 1980s under the stewardship of former Federal Open Market Committee (FOMC) chairmen Paul Volcker and Alan Greenspan. Sargent (2000) also argues that U.S. inflation returned to being low in the late 1980s and 1990s because of a combination of adaptive expectations and learning by a government with a simple view of the Phillips curve. Either explanation could be consistent with the TPC model failing to predict a secular decline in inflation.

Measuring the output gap with a Hodrick-Prescott filter could also be problematic. For example, many commentators argue that structural productivity growth increased in the second half of the 1990s in the United States, with a positive effect on potential output. To the extent that this break is not picked up by the HP filter, the output gap from the second half of the 1990s to the present could be overestimated. This measurement error could distort our estimates of the TPC and, indeed, lead to predicted inflation being upwardly biased.

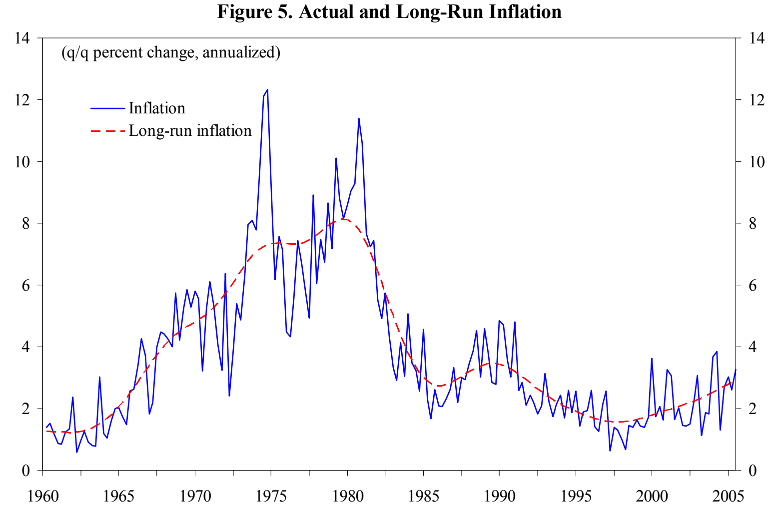
More fundamentally, output gap measures such as a HP filter are "high band-pass" filters and therefore remove secular components. Yet, the output gap is being used to explain the overall level of inflation in the TPC model. Such measures are better suited to explain movements in inflation relative to its long-run trend—namely the business-cycle component of inflation—rather than the secular trend itself. To investigate this issue further, and, more generally, the potential sources of misspecification in the TPC model, we now estimate the TPC model in the frequency domain.

C. Frequency Domain Analysis of TPC Model

Decomposing actual inflation into its "long-run" and "business-cycle/short-run" components may help to understand the overprediction problem. The previous analysis was carried out entirely in the time domain, in which both the short-run and long-run components of the data were used simultaneously to estimate the model. Because the "signal" of the business-cycle/short-run (hereon after referred to as "business-cycle") components tends to be overwhelmed by the size (and variance) of the long-run component, understanding the determinants of inflation over the business cycle using conventional time-series techniques can be problematic. By contrast, in this subsection we first decompose the inflation series (and its determinants) into independent long-run and business-cycle components and then estimate TPC models based on the filtered (i.e., business-cycle) series.

This paper uses recently developed frequency domain techniques to decompose inflation into its long-run and business-cycle components. In contrast to previous studies, the decomposition is carried out using a frequency domain technique that explicitly allows for nonstationary behavior, obviating the need to first make the series stationary, which often distorts the business-cycle component of the data (see Corbae, Ouliaris and Phillips, 2002, for a complete treatment and Appendix I for a simplified summary of the approach). To extract the business-cycle component, it is assumed that all variation in the data with cycles of less than 32 quarters (i.e., 8 years) belongs to the business-cycle component of the data. The long-run component is derived by simply subtracting the business-cycle component from the actual time-series variable.

Figure 5. Actual and Long-Run Inflation

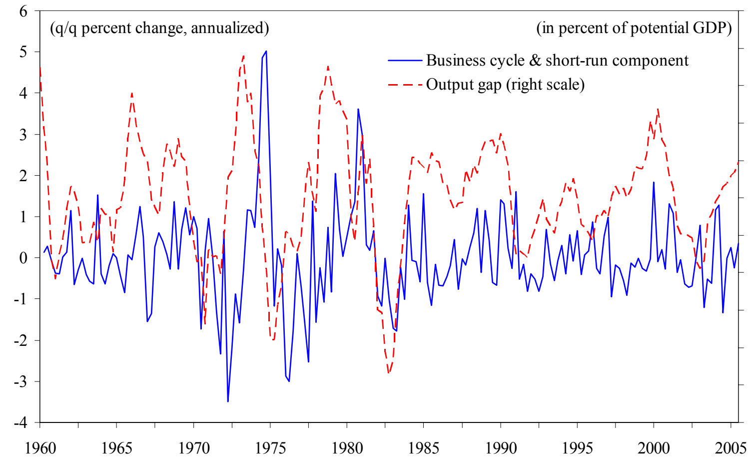


The behavior of long-run inflation relative to actual is shown in Figure 5. Clearly, the long-run component is simply a smoothed version of the actual inflation data, and can be interpreted as the (nonlinear) trend in the data. Figure 5 reveals a steep decline in trend inflation during 1982–98, after which it has been gradually increasing.[5] The volatility of actual inflation around its long-run path has also declined since 1982 (Figure 6). More specifically, the standard deviation of these business-cycle movements has halved since

[5] Calling turning points at the end of the time-series, however, is problematic using standard smoothing techniques. Clearly, the "upward" trend in the inflation series could be revised downward with a favorable outcome for inflation in the next period.

1982. Lastly, prior to 1982, movements in business-cycle component of inflation appear to lag movements in the output gap, but this relationship has weakened since 1982.

Figure 6. Business-Cycle and Short-Run Component of Inflation and the Output Gap



Estimating the conventional Phillips curve for 1961–2005 on the business-cycle components of inflation yields the following results:

$$\pi_t^{BC} = \underset{(0.0591)}{0.0011} + \underset{(2.5151)}{0.3322}\,\pi_{t-1}^{BC} + \underset{(0.1736)}{0.0140}\,\pi_{t-2}^{BC} + \underset{(0.5517)}{0.0421}\,\pi_{t-3}^{BC} + \underset{(2.1373)}{0.1663}\,\pi_{t-4}^{BC} + \underset{(4.8971)}{0.0482}(y_{t-1} - y_{t-1}^{*})$$
$$\overline{R}^2 = 0.2464 \quad \textit{Durbin-Watson} = 1.9885 \quad F_{(1,172)}(\sum_{1}^{4}\varphi_i = 1) = 11.8395 \ \ (\textit{p-value}: 0.0000) \tag{6}$$

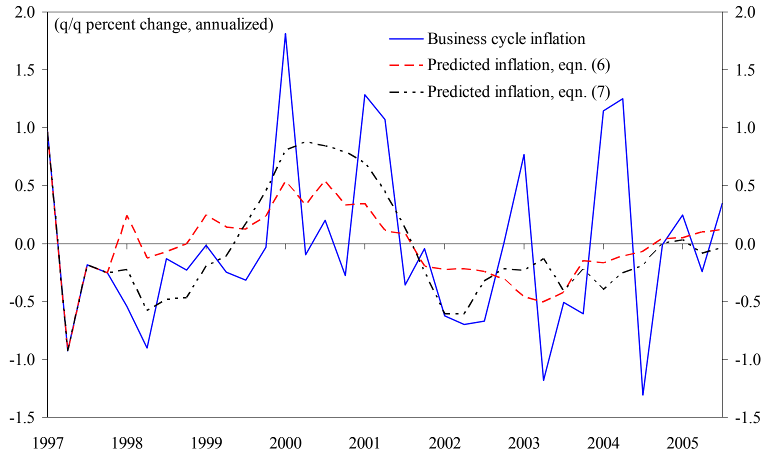
The significance of the lagged dependent variables over 1961–2005 confirms the persistence of inflation over the business-cycle. In contrast to the time-series results, however, the sum of the coefficients on lagged inflation implies the expected trade-off between inflation and the output gap over the business cycle.

Allowing for terms of trade effects improves the fit of the business-cycle model considerably. Terms of trade shocks relative to trend contribute positively to the business-cycle movements in inflation (equation (7)).

$$\pi_t^{BC} = \underset{(0.0833)}{0.00140} + \underset{(2.8671)}{0.2198}\,\pi_{t-1}^{BC} - \underset{(0.0881)}{0.0068}\,\pi_{t-2}^{BC} + \underset{(0.7006)}{0.0531}\,\pi_{t-3}^{BC} + \underset{(2.5164)}{0.1796}\,\pi_{t-4}^{BC} + \underset{(2.8248)}{0.0341}(y_{t-1} - y_{t-1}^{*})$$
$$+ \underset{(2.8181)}{0.0348}\,\Delta tot_{t-1}^{BC} + \underset{(2.3942)}{0.03134}\,\Delta tot_{t-2}^{BC} + \underset{(2.5738)}{0.03417}\,\Delta tot_{t-3}^{BC} + \underset{(0.9493)}{0.01266}\,\Delta tot_{t-4}^{BC} \tag{7}$$
$$\overline{R}^2 = 0.3735 \quad \textit{Durbin-Watson} = 1.9905 \quad F_{(1,168)}(\sum_{1}^{4}\varphi_i = 1) = 32.6398 \ \ (p - \textit{value}: 0.000)$$

Figure 7 suggests that the out-of-sample forecasting performance of equations (6) and (7) over 1998–2005 is reasonable. The forecasts confirm that these models are able to predict the more persistent business-cycle movements relative to trend. Interestingly, predictive ability over the business cycle improves considerably by allowing for terms of trade effects. Given the previous results for the TPC model, this suggests that terms of trade shocks are (a) mostly transitory in nature (i.e., relatively minor in the long run); and (b) more suited to explaining cyclical variations in inflation relative to trend.

Figure 7. Traditional Phillips Curve: Actual Versus Forecast Inflation over the Business Cycle



D. Taking Stock

TPC models appear to explain movements in inflation relative to trend, but have less success in explaining actual inflation owing to the existence of a declining secular trend. In particular, the model is unable to account for the secular decline in actual inflation since 1990—even after allowing for price-level shocks from the external sector.

III. New Keynesian Phillips Curve (NKPC)

NKPC models are based on microeconomic foundations that allow for forward-looking optimizing behavior. The NKPC approach is based on staggered price setting behavior. The underlying model assumes optimal price setting behavior by monopolistically competitive firms subject to the menu-costs imposed by the frequency of price adjustment (Calvo, 1983; Fischer, 1977; Taylor, 1980). Aggregating across firms leads to an aggregate Phillips curve that relates inflation to expectations of inflation and real marginal costs.

The NKPC model is robust to the Lucas critique, as it is derived from a microeconomic framework with optimizing firms and rational expectations. In contrast to TPC models, NKPC models use marginal cost to measure inflation pressures rather than the output gap. Expectations of future inflation capture the forward-looking nature of the inflation process, namely that optimizing firms will take into account future inflation expectations when they reset prices—as they may not have the opportunity to reset them again for a while.

A. Hybrid Model

We follow Galí and Gertler (1999) and allow for backward-looking behavior in our baseline NKPC model. Previous studies have found that allowing firms to be sclerotic when they reset prices—in other words allowing for the impact of lagged inflation—generates a better fit of the model. This "hybrid version" of the NKPC model relates inflation to past and expected future inflation, and changes in real marginal cost relative to its steady state level:

$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f E\{\pi_{t+1}\} + \lambda mc_t \tag{8}$$

where λ is a slope coefficient that depends on the parameters of the underlying price optimization model, especially the degree of price rigidity. In equation (8), γ_b and γ_f measure the degree of backward and forward-looking behavior respectively, with some researchers imposing the constraint that $\gamma_b + \gamma_f = 1$. Moreover, γ_b / γ_f measures the degree of forward-looking behavior in the model.

The link between the hybrid NKPC model and the traditional Phillips curve becomes apparent given the relationship between marginal cost and the output gap. Rotemberg and Woodford (1999) show that under specific restrictions on technology and labor market structure, real marginal cost moves in proportion to the output gap. In this case, the NKPC model becomes:

$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f E\{\pi_{t+1}\} + \tau(y_t - y_t^*) \tag{9}$$

where, as before, $y_t - y_t^*$ is the log deviation of real GDP from its long run trend. In contrast to the TPC, inflation depends on expectations regarding future movements in the output gap, and, hence, it should lead movements in the output gap.[6] However, as Galí, Gertler, and López-Salido (2001) demonstrate, the simple link between marginal cost and the output gap is invalid if there are frictions in the labor market.

B. Basic Econometric Specification

In line with most earlier work, we use the labor share of GDP as a proxy for marginal cost. A close relationship between labor share and marginal cost holds when the firm-level

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[6] Preliminary estimates (unreported) provided little support for this model. Consistent with the findings of Galí and Gertler (1999), the parameter estimates for τ were of the wrong sign (negative) and statistically insignificant. Equation (9) is not used for this reason.

production function is Cobb-Douglas. Other possibilities would have been to allow for labor or capital adjustment costs, and to consider a constant elasticity substitution (CES) production function. As Balakrishnan and López-Salido (2002) document for the United Kingdom, however, the time-series properties of real marginal cost measures tend to be very similar in the different cases. Moreover, we focus on the reduced form of the NKPC model (equation (8)), as opposed to the structural form. The latter would make λ, γ_b and γ_f functions of deeper parameters such as the frequency of price adjustment, the fraction of firms setting prices optimally, and the subjective discount factor. To the extent that estimates of these deeper parameters are not a focus of this paper, the choice of production function or whether capital is firm specific (see Woodford, 2005) should not be material.

We estimate the NKPC using generalized methods of moments (GMM). While this approach has come under criticism recently (Rudd and Whelan, 2005; and Lindé, 2005), Galí, Gertler, and López-Salido (2005) argue that the results emanating from a GMM procedure are consistent with those of other procedures such as nonlinear instrumental variables and full information maximum likelihood. Against this background, we use GMM, based on the following orthogonality condition:

$$E_{t-1}\{(\pi_t - \gamma_f \pi_{t+1} - \gamma_b \pi_{t-1} - \lambda mc_t) z_{t-1}\} = 0 \qquad (10)$$

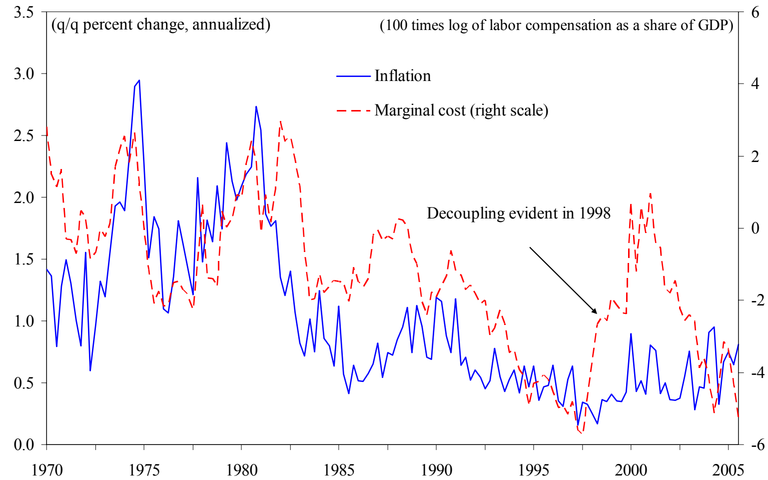
To avoid potential estimation bias that is common in small samples when there are too many over-identifying restrictions, we choose our instruments parsimoniously, with our vector of instruments, z_t, including: four lags of inflation and two lags of marginal cost, HP filtered real output and nominal wage inflation.

C. Time Domain Regression Estimates of NKPC Model

The relationship between inflation and marginal cost is generally positive, but decouples in the late 1990s (Figure 8). This suggests that other factors—not accounted for in the basic NKPC model—have kept inflation low since then.[7]

[7] Our measure of the labor share is based on nonfarm labor compensation, which includes stock options. Adjusting for stock options—based on a sample of some of the largest S&P firms which accounted for about 70 percent of market value from 1995–2002—makes little difference to the decoupling.

Figure 8. Inflation and Marginal Cost (Labor Share)



Over 1970–2005, behavior is mostly forward-looking, with a minor role for marginal cost movements. The baseline inflation equation is given by:

$$\pi_t = \underset{(1.3965)}{0.0443} + \underset{(7.9612)}{0.6377}\,\pi_{t+1} + \underset{(4.1759)}{0.3301}\,\pi_{t-1} + \underset{(1.7853)}{0.0127}\,mc_t \qquad (11)$$
$$\overline{R}^2 = 0.8254 \quad Durbin\text{-}Watson = 3.0567$$

The parameter estimates imply that movements in marginal cost have a positive (and marginally statistically significant) impact on inflation. Backward-looking behavior is also evident, although expectations concerning inflation play a far greater role in explaining actual inflation—the estimates imply that inflationary expectations account for over 75 percent of the explanatory power of the model. The dominance of forward-looking behavior is consistent with previous results for both the United States and euro area (e.g., Galí and Gertler (1999); Galí, Gertler, and López-Salido (2001); and Tillmann (2005)). The contribution of the marginal cost term to the explanatory power of the model is less than 4 percent—an issue we return to in subsection III.D.

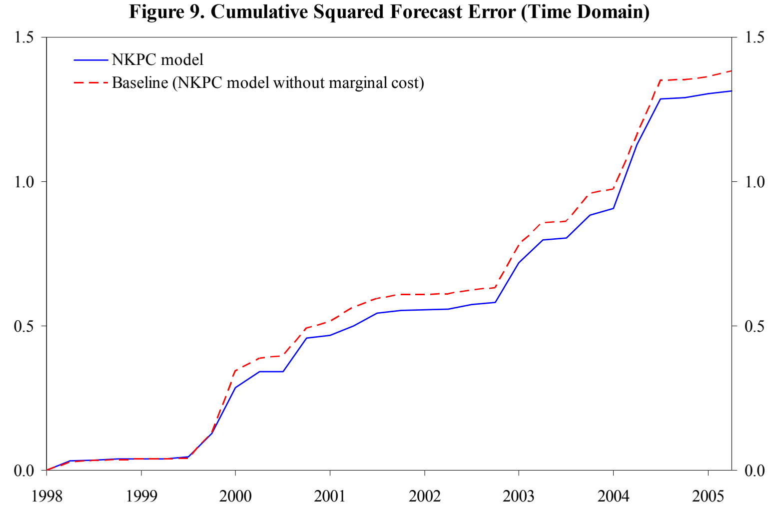
Reflecting the decoupling after 1997, the relationship between inflation and marginal cost is slightly stronger over 1970–97 (equation (12)). These results confirm the findings of Galí and Gertler (who estimated the NKPC over the same time period) with the marginal cost term significant and contributing approximately 7 percent to the overall explanatory power of the model.

$$\pi_t = \underset{(1.8128)}{0.0774} + \underset{(7.1613)}{0.6355}\,\pi_{t+1} + \underset{(3.6183)}{0.3117}\,\pi_{t-1} + \underset{(1.8907)}{0.0173}\,mc_t \qquad (12)$$
$$\overline{R}^2 = 0.8128 \; Durbin - Watson = 3.0283$$

D. Performance of Basic Model and Frequency Domain Results

Perfect foresight forecasts reaffirm that marginal cost plays little role in explaining inflation in the time domain. The NKPC has current inflation dependent on expectations of inflation, which given the need for expectations to be model-consistent, makes using the model for forecasting problematic. Nevertheless, one can assume agents have perfect foresight (i.e. inflation expectations are realized) and use this path to test the stability of the model. We estimate the model with and without marginal cost over 1970–97. Comparing the mean squared error (MSE) of the perfect foresight forecasts from 1998–2005 (Figure 9) suggests a limited role for marginal costs.

Figure 9. Cumulative Squared Forecast Error (Time Domain)



The limited role for marginal cost might be due to the fact that the model is better suited to explaining inflation over the business cycle. Two considerations come to mind:

- *Positive trend inflation*. Theoretically, the NKPC seeks to explain inflation deviations from trend and in effect assumes that trend inflation is zero. Woodford (2003), and Bakshi, Khan, and Rudolf (2004) argue, however, that the hybrid NKPC is a good proxy for a general class of sticky price models of actual inflation, provided that trend inflation is low.

- *Inflation leads and lags*. Consistent with the previous argument, because the NKPC model explains actual inflation using its own lead and lag, it technically falls into the class of high band-pass filters (see Baxter and King, 1999) that by construction moderate the trend. This means that the NKPC model is better suited to predicting business-cycle/short-run movements in inflation (i.e., movements in inflation relative to trend).

Simply introducing trend inflation in the NKPC model is unlikely to increase the limited role for marginal costs. Cogley and Sbordone (2005) extend the NKPC to explicitly allow for positive trend inflation. They argue that the link between current marginal cost and inflation is weakened as trend inflation increases, and that the influence of forward-looking terms is enhanced. This follows from the fact that as trend inflation accelerates the rate at which a firm's relative price is eroded increases when it lacks opportunity to reset prices. This channel goes against the U.S. evidence, however, which shows that inflation and marginal costs decoupled when trend inflation was relatively low.

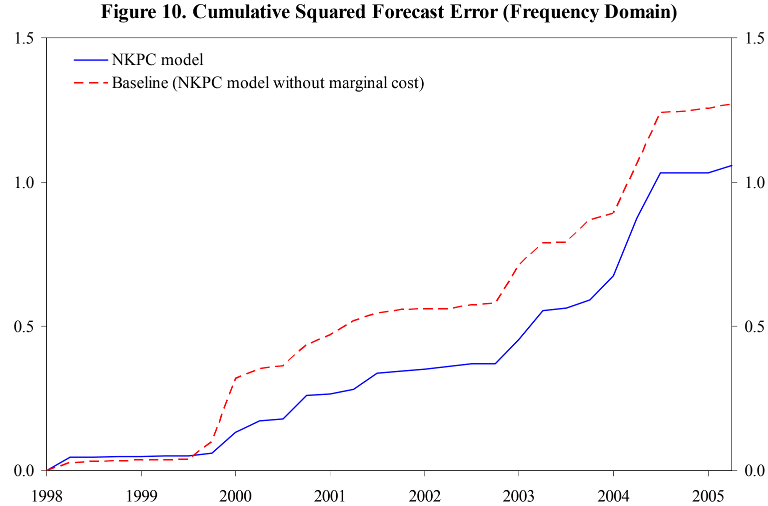
Estimating the NKPC using business-cycle data gives an increasing role to marginal cost relative to the time domain model, and a reduced role for inflation expectations. Equations (13) and (14) show the results of estimating the NKPC model using business-cycle data for 1970–2005 and 1970–97 respectively, following the methodology outlined in subsection II.C.

$$\begin{aligned} \pi_t &= \underset{(0.3287)}{-0.0043} + \underset{(2.2207)}{0.3154}\,\pi^{BC}_{t+1} + \underset{(4.3634)}{0.2973}\,\pi^{BC}_{t-1} + \underset{(1.8697)}{0.0762}\,mc_t \\ \overline{R}^2 &= 0.3049 \;\; Durbin-Watson = 2.6357 \end{aligned} \tag{13}$$

$$\begin{aligned} \pi_t &= \underset{(0.2262)}{-0.0038} + \underset{(2.3245)}{0.3306}\,\pi^{BC}_{t+1} + \underset{(4.4437)}{0.3006}\,\pi^{BC}_{t-1} + \underset{(1.8157)}{0.0817}\,mc_t \\ \overline{R}^2 &= 0.3270 \;\; Durbin-Watson = 2.6688 \end{aligned} \tag{14}$$

The magnitude of the expected inflation term declines significantly, but still remains significant, as does the coefficient on lagged inflation. This result reaffirms the finding using TPC models that inflation is highly persistent over the business-cycle frequencies. Notably, the coefficient on marginal cost increases significantly relative to the time-domain model, and this is reflected in the MSE of the perfect foresight forecast including marginal cost being significantly lower than the one just based on the lead and lag of inflation (Figure 10).

Figure 10. Cumulative Squared Forecast Error (Frequency Domain)



E. Extended NKPC Model

Next, we extend the basic NKPC model to take account of the impact of increasing globalization. Many analysts have argued that increasing real and financial globalization are magnifying the effect of external shocks on the domestic economy—even on a largely closed economy like the United States (IMF 2005). One would also expect the large increases in commodity prices in recent years to have impacted on inflation. To capture both of these effects, we add imported intermediate goods to our firm-level production function, and allow the desired markup to vary over time.

We allow the desired markup to vary with the cycle and external competition. The desired markup will depend on the demand elasticity facing the firm, which, in turn, depends on the degree of competition faced by the firm in the product market. Following Rotemberg and Woodford (1999), it is easy show that the NKPC with variable desired markups becomes:

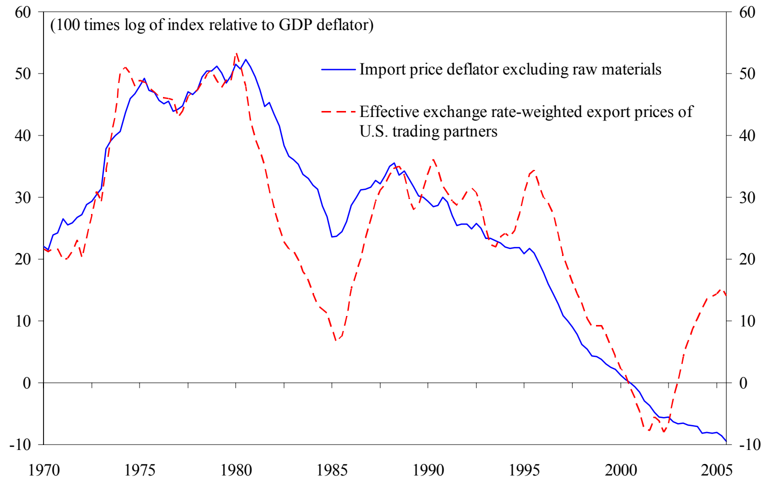
$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f E\{\pi_{t+1}\} + \lambda mc_t + \hat{\mu}_t \tag{15}$$

where $\hat{\mu}$ is the (log) desired markup relative to its steady state. We consider two channels:

- *Cyclical factors (output gap)*. Rotemberg and Woodford (1999) argue that the impact of the business cycle on desired markups can be positive or negative depending on whether the main factor is a varying elasticity of demand, customer markets, implicit collusion, or variable entry. For example, regarding prices wars, some models predict that they are more likely in booms, and others in slumps. We do not take a stand on this, and simply assume that the output gap can affect the desired markup and include it in the base specification.

- *External competition*. Following Batini, Jackson, and Nickell (2005) we also proxy for the degree (or threat) effect of foreign competition. Batini, Jackson, and Nickell (2005) use the effective exchange weighted export prices of the main trading partners of the country concerned, which was the United Kingdom in their study. We find, however, that the export price index is a poor proxy of foreign competition as it fails to control the composition of exports reaching the country concerned, in our case the United States. Indeed, we prefer the U.S import price deflator net of the effect of raw material imports, since it directly captures noncommodity imports into the United States. Figure 11 plots the two series and illustrates a marked difference, with the export price index having increased since 2002 and the import price deflator continuing its decline, the latter suggesting that external competition has strengthened in recent years.

Figure 11. Competition Indices



We also make production dependent on imported intermediate goods. Again, we follow Batini, Jackson, and Nickell (2005) and assume that gross output requires imported intermediate goods in the following way:

$$M = m(Y_g)Y_g \quad \text{with } m' > 0 \tag{16}$$

where M is the amount of imported intermediate goods needed and Y_g is gross output. This results in marginal cost (relative to its steady state) being a function of the price of imported intermediate goods—which we proxy by the IMF total commodity price index for the United States—as well as the labor share (relative to its steady state), $\hat{s}$, namely:

$$mc = \hat{s} + \beta_m(p_m - p) \tag{17}$$

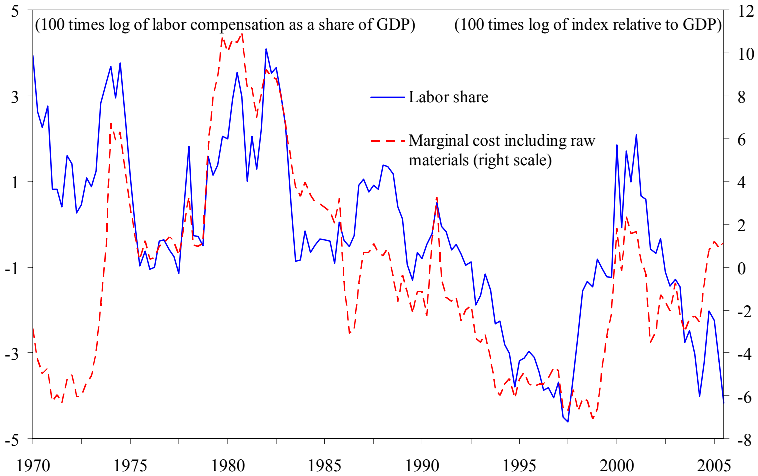
where $p_t^m - p_t$ is the (log) price of imported materials relative to domestic goods. Our extended NKPC model may therefore be represented as

$$\begin{aligned} \pi_t &= \gamma_b \pi_{t-1} + \gamma_f E\{\pi_{t+1}\} + \lambda mc_t + \hat{\mu}_t, \\ \hat{\mu}_t &= \alpha_0 + \beta_x Comp_X + \beta_{y-y^*}(y_t - y_t^*), \\ mc_t &= \alpha_1 + \hat{s}_t + \beta_m(p_t^m - p_t) \end{aligned} \tag{18}$$

where $Comp_X$ captures external competitive pressures.

The inclusion of imported intermediate goods mitigates the decoupling but does not eliminate it. Figure 12 shows that marginal cost with and without imported intermediate goods have broadly similar trends. However marginal cost including imported intermediate goods increases less from 1997–2001 and actually starts increasing again in 2003, with the latter probably reflecting the substantial runup in commodity prices.

Figure 12. Marginal Cost Measures



Against this background, as for the basic NKPC model, the extended model has a reasonable fit through 1997 but subsequently breaks down (Table 1, Rows 1 and 2). Marginal cost is insignificant over 1970–2005, but is strongly significant through 1997. Indeed, the size of the coefficient on marginal cost is much higher over 1970–97 than for the whole sample period. The forward-looking nature of the inflation process is still clear, though less so than in the basic NKPC model (i.e., equation (11)). Neither the output gap nor external competition measures are significant over 1970–2005.

Table 1. Empirical Results for New Keynesian Phillips Curve:
Time Domain and Business-Cycle Frequencies

Model	γ_f (Future Inflation)	γ_b (Lagged Inflation)	λ (Marginal Cost)	β_m (Relative Import Price Deflator)	β_{y-y^*} (Output Gap)	β_x (External Competition)
Time domain						
1970:1–2005:4	0.5369 (5.4745)	0.3586 (3.9283)	0.0109 (1.5516)	0.0862 (1.3650)	0.0105 (1.2512)	$0.8023e-03$ (0.9161)
1970:1–1997:4	0.4987 (5.2788)	0.3989 (4.1569)	0.0608 (3.4866)	0.0265 (3.0897)	0.0323 (2.3732)	$-0.1131e-01$ (1.7558)
Business-Cycle						
1970:1–2005:4	−0.4140 (1.70306)	0.0818 (1.0543)	0.1404 (2.1116)	0.0270 (1.6265)	0.0213 (1.0802)	$0.7890e-01$ (2.9938)
1970:1–1997:4	−0.4901 (3.0462)	0.1004 (1.2027)	0.0754 (1.7243)	0.0541 (1.2050)	0.0555 (4.2175)	$0.1013e-03$ (4.8163)

Over business-cycle frequencies, however, the importance of inflation expectations declines, while that of external competition and imported intermediate goods increases (Table 1, Row 3). Following the methodology of subsection II.C, we note that marginal cost is significant and its impact is much higher than in the basic NKPC models (both frequency and time domain versions). Moreover, the inflation lead and lag terms are, respectively, wrongly signed and statistically insignificant. Relative to the time domain estimates for the extended model, the size of the estimated business-cycle coefficients on both the external competition measure (highly significant) and the imported intermediate goods variable (marginally significant) increase substantially and are correctly signed. Moreover, the contribution of the external competition variable to the overall fit of the model exceeds 40 percent—compared to less than 5 percent for the time-domain model. In line with the stylized facts presented in Figure 5 concerning the volatility of the business-cycle component, these findings suggests that increased external competition has moderated the business-cycle component of inflation after 1982.

IV. Conclusions and Policy Implications

TPC models suggest that the decline in inflation since 1990 has more to do with the behavior of the long-run (or secular) trend than the business-cycle component. Like earlier models, the TPC fails to account for the recent trend decline in inflation, although incorporating external shocks—proxied by movements in the terms of trade—reduces the degree of overprediction. By contrast, the TPC model produces acceptable out-of-sample forecasts for the business-cycle component of inflation. When combined with the fact that inflation has also become less variable relative to trend since the late 1980s, these results suggest that the decline in inflation is indeed a structural—as opposed to cyclical—phenomenon.

While the basic NKPC model also fails to fully explain the decline in inflation, the analysis suggests that the link between marginal cost and inflation has weakened in recent years. NKPC models were developed to address the shortcomings of TPC models, particularly their backward-looking nature and sensitivity to policy regime shifts. NKPC models of inflation allow for forward-looking inflation expectations and focus on marginal cost as opposed to measures of the output gap, and we find that the relationship between inflation and marginal cost breaks down in the late 1990s—around the same time that TPC models starts to overpredict inflation.

As for the TPC model, external variables improve the fit of the NKPC model, supporting the hypothesis that trade and global factor markets have helped to moderate inflation, although mainly over the business cycle. The basic NKPC model was extended by introducing imported intermediate goods and allowing desired price markups to vary with the business cycle and degree of competition. This extension improves the overall performance of the model, although the impact and significance of external variables is evident mainly in the cyclical component.

Overall, we find that excess demand, external shocks, and competitive pressures linked to globalization help to explain cyclical movements in inflation relative to trend. The empirical analysis also suggests that these variables are unlikely to keep inflation low on a permanent basis, since they fail to resolve the overprediction problem. This finding is consistent with the published views of former Fed Chairman Alan Greenspan, who has argued that as globalization effects wane, policymakers will need to work harder to keep inflation pressures contained.

What explains the secular decline in inflation? Notwithstanding the absence of satisfactory proxies for monetary policy credibility, the long-run decline in actual inflation could reflect improved U.S. monetary policy credibility and the adoption of an implicit inflation-targeting regime. Indeed, augmenting time-domain TPC models with a satisfactory proxy for improved monetary policy credibility should improve the out-of-sample prediction of the TPC model. Higher credibility appears to have not only contributed to the trend decline in inflation but also lowered the variability of inflation around its long-run path. Pinning down the structural and cyclical implications of monetary policy credibility will be an important project for future research.

Extracting Business Cycles from Nonstationary Data

If one accepts the Burns and Mitchell (1946) definition of the business cycle as fluctuations in the level of a series within a specified range of periodicities, then the ideal filter is simply a band-pass filter that extracts components of the time series with periodic fluctuations between 6 and 32 quarters (see Baxter and King (1999)).[8] It can be shown that the exact band-pass filter is a double-sided moving average of infinite order, and with known weights. It follows that if we want to estimate the filter starting from the time domain, an approximation to the correct result is needed.

Many approximations have been suggested in the literature, with perhaps the Hodrick and Prescott (1980) and Baxter and King (1999) filters being the most popular. In this appendix, we outline an alternative, *frequency domain,* procedure for approximating the ideal band-pass filter, originally suggested in Corbae and Ouliaris (2006), which overcomes some of the shortcomings of the Hodrick and Prescott (1980) and Baxter and King (1999) time domain based filters (see below).

Assume that x_t $(t = 1, \ldots, n)$ is an observable time series generated by:

$$x_t = \Pi_2' z_t + \tilde{x}_t, \qquad \text{(A1)}$$

where z_t is a p+1-dimensional deterministic sequence and $\tilde{x}_t$ is a zero mean time series. The series x_t therefore has both a deterministic component involving the sequence z_t and a stochastic (latent) component $\tilde{x}_t$. In developing their approach to estimating ideal band-pass filters, Corbae and Ouliaris (2006) make the following assumptions about z_t and $\tilde{x}_t$.

Assumption 1

$z_t = (1, t, \ldots, t^p)'$ *is a* p^{th} *order polynomial in time.*

Assumption 2

$\tilde{x}_t$ *is an integrated process of order one (I(1) process) satisfying* $\Delta \tilde{x}_t = v_t$, *initialized at* $t = 0$ *by any* $O_p(1)$ *random variable. We assume that* v_t *has a Wold representation*

$v_t = \sum_{j=0}^{\infty} c_j \xi_{t-j}$ *where*

$\xi_t = iid\ (0, \sigma^2)$ *with finite fourth moments and coefficients* c_j *satisfying* $\sum_{j=0}^{\infty} j^{1/2} |c_j| < \infty$. *The spectral density of* v_t *is* $f_{vv}(\lambda) > 0$, $\forall\lambda$.

[8] Researchers of the business-cycle, including Burns and Mitchell, do not necessarily accept this definition of the growth cycle. For example, Harding and Pagan (2002) would include movements in trend (or zero frequency elements) as a fundamental part of cyclical movements.

Assumption 2 suffices for partial sums of v_t to satisfy the functional law $n^{-1/2} \sum_{t=1}^{[nr]} v_t \xrightarrow{d} B(r) = BM(\sigma^2)$, a univariate Brownian motion with variance $\sigma^2 = 2\pi f_{vv}(0)$ (e.g., Phillips and Solo (1992), theorem 3.4), and where $\xrightarrow{d}$ is used to denote weak convergence of the associated probability measures as the sample size $n \to \infty$. We now state the result that motivates the new filtering procedure.

Lemma B (Corbae, Ouliaris, and Phillips (2002)) *Let* $\tilde{x}_t$ *be an* I(1) *process satisfying Assumption 2. Then, the discrete Fourier transform of* $\tilde{x}_t$ *for* $\lambda_s \neq 0$ *is given by*

$$w_{\tilde{x}}(\lambda_s) = \frac{1}{1-e^{i\lambda_s}} w_v(\lambda_s) - \frac{e^{i\lambda_s}}{1-e^{i\lambda_s}} \frac{[\tilde{x}_n - \tilde{x}_0]}{n^{1/2}} \tag{A2}$$

where the discrete Fourier transform (dft) of $\{a_t; t = 1, \ldots, n\}$ *is written* $w_a(\lambda) = \frac{1}{\sqrt{n}} \sum_{t=1}^{n} a_t e^{i\lambda_t}$, *and* $\{\lambda_s = \frac{2\pi s}{n}, s = 0, 1, \ldots, n-1\}$ *are the fundamental frequencies.*

Equation (A2) shows that the discrete Fourier transforms of an I(1) process are not asymptotically independent across fundamental frequencies. They are actually frequency-wise dependent by virtue of the component $n^{-1/2}\tilde{x}_n$, which produces a common leakage into all frequencies $\lambda_s \neq 0$, even in the limit as $n \to \infty$. Corbae, Ouliaris, and Phillips (2002) also show that the leakage is still manifest when the data are first detrended in the time domain. These results on leakage show that in the presence of I(1) variables, any frequency domain estimate of the "cyclical" component of a time series (e.g., real GDP) will be badly distorted.

Corbae and Ouliaris (2006) suggest a simple "frequency domain fix" to this problem, which is derived from equation (A2). Note that the second expression in equation (A2) can be rewritten using

$$w_{\left(\frac{t}{n}\right)}(\lambda_s) = \frac{-1}{\sqrt{n}} \left(\frac{e^{i\lambda_s}}{1-e^{i\lambda_s}} \right)$$

by Lemma B of Corbae, Ouliaris, and Phillips (2002). Thus, even for the case where there is no deterministic trend in equation (A1), it is clear from the second term in equation (A2), which is a deterministic trend in the frequency domain with a random coefficient $[\tilde{x}_n - \tilde{x}_0]$, that the leakage from the low frequency can be removed by simply detrending in the frequency domain, leaving an asymptotically unbiased estimate of the first term $\frac{1}{1-e^{i\lambda_s}} w_v(\lambda_s)$ over the nonzero frequencies. We recommend that this detrending be done *before* the relevant business-cycle frequencies are identified, as this will maximize the number of frequency domain terms available to estimate $\left[\tilde{x}_n - \tilde{x}_0\right]$.[9] An indicator function can then be applied to the unbiased estimate of $\frac{1}{1-e^{i\lambda_s}} w_v(\lambda_s)$ to annihilate (or zero out) the non-business-cycle frequencies. For example, for the Burns and Mitchell (1946) definition of the business cycle, all frequencies outside the [6, 32] quarter range would be set to zero, while those within the range would be kept at their original values.[10] The filtered series is then obtained by applying the inverse Fourier transform to the result. It can be shown that the filtered series will be a $\sqrt{n}$ consistent estimate of the true business cycle over the included frequencies (see Corbae and Ouliaris (2006)).

[9] Our recommendation follows from the fact the true coefficient $\left[\tilde{x}_n - \tilde{x}_0\right]$ in equation (A2) does not vary with frequency.

[10] The indicator function would have a value of unity for each frequency that needed to be included in the filter, and zero for each frequency that needed to be excluded from the filter. For example, for the classical business cycle, the indicator function would have a value of unity for all fundamental frequencies that fall in the range $\left[\frac{\pi}{16}, \frac{\pi}{3}\right]$. GAUSS code for computing the frequency domain (FD) filter is available from the authors on request.

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