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Are prices countercyclical? Evidence from the G-7

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Abstract

This paper re-examines the cyclical behavior of prices using postwar quarterly data for the G-7. We confirm recent evidence that the price level is countercyclical. However, we find strong evidence that the inflation rate is procyclical in our sample. Our results show the importance of making a clear distinction between inflation and the cyclical component of the price level when reporting and interpreting stylized facts regarding business cycles.

Key words: Business cycles; Inflation; Prices

JEL classification: E31; E32

1. Introduction

The objective of business cycle theory is to explain a set of stylized facts within a consistent framework. A key stylized fact that has served as a fundamental datum in the construction of a large class of business cycle models is the procyclical behavior of prices. Recent papers by Kydland and Prescott (1990), Cooley and Ohanian (1991), Backus and Kehoe (1992), and Smith (1992) have challenged the empirical basis for this stylized fact. These papers show that the

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cyclical components of prices and output are negatively correlated and interpret this finding as a falsification of the conventional wisdom that prices are procyclical.

Since the price level and the level of output are both nonstationary series, the procedure used in these papers is to independently detrend both the price and output series using a common filter to remove the nonstationary components and to examine the correlations between the stationary residuals. In the most comprehensive of the recent studies of cyclical price variation for the United States, Cooley and Ohanian (1991), using both postwar and historical data, compute three alternative cyclical measures of output and prices: the first differences of the series, deviations of the series from a constant and a time trend, and filtered values of the series using the Hodrick–Prescott filter. An examination of the cross-correlations of the resulting series suggests that procyclical price movements have not been a stable feature of business cycles in the United States. In particular, for the postwar period the cross-correlations of the series are typically significantly negative at most lags and leads using any of the three cyclical measures. This is prima facie evidence for countercyclical price behavior. Similar evidence is presented by Backus and Kehoe (1992) and Smith (1992) using annual data for a larger set of industrialized economies.

However, the traditional focus of business cycle models, including Keynesian demand-driven models as well as rational expectations models along the lines of Lucas (1972, 1976), has been the inflation rate rather than the price level. For instance, in motivating his theoretical model, Lucas (1976) points out that 'the fact that nominal prices and wages tend to rise more rapidly at the peak of the business cycle than they do in the trough has been well recognized from the time when the cycle was first perceived as a distinct phenomenon'. Although one might expect the price level and the inflation rate to exhibit similar cyclical behavior, there is no compelling theoretical reason for this to be true.

The empirical resolution of this issue also has implications for the sources of business cycles. A central question in explaining fluctuations in aggregate economic activity is whether short-run deviations of output from a longer-term (deterministic or stochastic) trend are attributable primarily to movements in, or shocks to, demand or supply. If temporary movements of output result primarily from shocks to demand, prices would be expected to be procyclical; if they result from shocks to supply, prices would be expected to be countercyclical. It is widely perceived that temporary movements in output are associated with shocks to demand, while longer-term movements are associated with movements in supply (see, e.g., Blanchard and Quah, 1989). However, the countercyclical variation of prices documented in the papers cited above suggests that even temporary movements in output may be due to supply disturbances. This is consistent with the view that supply-driven models of the business cycle, including the class of recently developed real business cycle models, may be more accurate representations of reality than conventional demand-driven models.

This evidence also counters the criticism that Mankiw (1989) has levelled at real business cycle models on the grounds that they imply that prices are not procyclical.

The objective of this paper is to provide a set of stylized facts for the main industrialized economies that clearly differentiates between the cyclical behavior of inflation and the price level. The current debate between proponents of alternative classes of business cycle models suggests that, in discriminating between theoretical models on the basis of stylized facts, this distinction is potentially very important. This distinction is also important for empirical models of the cycle in motivating identifying restrictions that are based on the cyclical behavior of prices.

At a purely empirical level, this paper's contribution is to ascertain if the countercyclical variation of the price level in the U.S. and other industrialized economies carries over to countercyclical variation in inflation. Confirming the results of the papers cited above, we find that, in postwar quarterly data for the U.S. as well as the rest of the G-7, when the levels of prices and output are subjected to the same stationarity-inducing transformation, the residuals of the two series are negatively correlated. The quantitative results are affected by the choice of the detrending procedure but the hypothesis of countercyclical price behavior is supported in most cases.

The key finding of this paper is that the inflation rate is generally positively correlated with various measures of the cyclical component of output. Although again the results are sensitive to the choice of the procedure for detrending output, this conclusion holds in most cases for the G-7. The positive correlation of inflation with the business cycle suggests to us that demand-driven models of the business cycle are not necessarily falsified by the countercyclical behavior of the price level. The striking contrast between the cyclical behavior of the (detrended) price level and inflation indicates the need for caution in examining and interpreting price cyclicality. Our findings highlight the importance of defining clearly whether the inflation rate or the cyclical component of the price level is the appropriate variable for analysis in a particular model. In this paper, however, we do not attempt to reconcile the countercyclical variation in the price level with procyclical variation in inflation.[1]

We also examine the cyclical behavior of prices and inflation using the unemployment rate as an alternative indicator of the cycle. This set of stylized facts is clearly relevant for the Phillips curve literature, although in this paper we do not interpret our results in that context. The HP-filtered unemployment rate is shown to be negatively correlated with inflation over our sample period for

[1]Chadha and Prasad (1993) have examined this issue more closely for the United States. That paper offers a potential explanation for these results on the basis of the detrending techniques used for output and the price level.

the G-7, confirming that inflation varies procyclically. However, the evidence on the countercyclical behavior of the price level is much less robust using the unemployment rate as an indicator of the cycle.

The next section of the paper presents empirical correlations of prices and inflation with cyclical output. Section 3 reports correlations of prices and inflation with the unemployment rate. Section 4 summarizes the main results and concludes.

2. Empirical results

This section first presents correlations between the levels of prices and output in the G-7, with both series rendered stationary using the same stationarity-inducing transformation. Next, we examine the correlations between inflation and various measures of cyclical output. The output series used in this study is quarterly real GDP (or, in the cases of Germany and Japan, real GNP). The price series employed is the implicit deflator corresponding to the concept of output used for a particular country. The price and output series were obtained from the IMF's International Financial Statistics tape.

2.1. Correlations between prices and cyclical output

Aggregate output and the price level are both nonstationary for the G-7 countries in the postwar period. A substantial literature has developed that studies the nature of nonstationarity in industrial country output (e.g., Cogley, 1990; Campbell and Mankiw, 1989).[2] The critical issue is whether output should be modelled as being stationary around a deterministic trend or a stochastic trend. The characterization of the form of nonstationarity is important since this determines the procedure for estimating the trend component of output and, thereby, for deriving the appropriate measure of the cyclical component. Although the nature of the nonstationarity of the price level has not been investigated as extensively, the concerns are similar (see, e.g., Balke and Fomby, 1991). Rather than restrict ourselves to a particular technique for detrending output and prices, we prefer to take an eclectic approach and evaluate the robustness of our results across a variety of detrending procedures.

In examining price–output correlations, we use four techniques to transform prices and output into stationary series.[3] First, we detrend output and prices

[2]A large body of literature (e.g., Cochrane, 1988; Perron, 1989; Christiano and Eichenbaum, 1990; Rudebusch, 1993) has been devoted to analyzing this issue for the U.S., although it appears that no conclusive resolution has yet been reached about how best to characterize the nonstationary component of U.S. output.

[3]All series are first transformed into logarithms.

using a deterministic linear trend. Second, we again detrend prices using a linear trend but, along the lines suggested by the work of Perron (1989, 1990), we detrend output using a segmented linear trend, with a break in the level and slope of the trend in 1974:1. This period corresponds to the first oil shock and the subsequent productivity slowdown in most of the G-7. Third, we use the Hodrick–Prescott (HP) filter to detrend the price and output series.[4] Fourth, we take the first differences of both prices and output.

The first panel of Table 1 reports the correlations between detrended prices and up to four lags and leads of detrended output, with both series detrended using a simple linear trend. The standard errors for the correlation coefficients were very similar at various leads and lags.[5] Hence, to conserve space, we present only the maximum standard error at any lag or lead. In this panel, the correlations for Germany and Japan are strongly positive and those for France are positive at the lags. For all other countries, the correlations are strongly negative, indicating countercyclical price behavior.

The above results are subject to the criticism that linear detrending of output in G-7 countries may be inappropriate as the output series in these countries may not be trend-stationary and may in fact be difference-stationary (Campbell and Mankiw, 1989). However, Perron (1990) argues that the hypothesis that output in the G-7 has a unit root can be rejected in favor of the alternative hypothesis that output is stationary around a linear trend with one break in the slope of the trend around the time of the first OPEC oil shock (1973–74).[6] Following Perron, we detrend output using a segmented linear trend, with a break in the trend in 1974:1.

The correlations between this measure of cyclical output and prices detrended as before, using a linear trend, are reported in the second panel of Table 1. Since there is no clear evidence regarding shifts in the trend price level for the G-7, we do not allow for a break in the estimated trend for the price level. The price–output correlations now turn strongly negative for France and remain strongly negative for Canada, Italy, the U.K., and the U.S. However, for Germany and Japan, the correlations remain strongly positive.

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[4]The Hodrick–Prescott filter involves solving the following minimization problem:

$$\min_{\{q_t\}} \frac{1}{T}\sum_{t=1}^{T}(y_t - q_t)^2 + \frac{\lambda}{T}\sum_{t=2}^{T-1}[(q_{t+1} - q_t) - (q_t - q_{t-1})]^2,$$

where y_t is the original series, q_t is the trend or growth component, and $y_t - q_t$ is the residual. In our computations, we set $\lambda = 1600$ as suggested by Prescott (1986) for quarterly data. See Cogley and Nason (1991) and King and Rebelo (1993) for an analysis of the properties of the Hodrick–Prescott filter.

[5]Note that these are approximate standard errors, computed under the null hypothesis that the true correlation coefficient is zero.

[6]This argument is disputed by Banerjee, Lumsdaine, and Stock (1992).

The third panel of Table 1 presents correlations between HP-filtered prices and output. A few of the lagged correlations for Canada, Germany, Italy, and Japan are positive in this panel. For the remaining countries, most of the correlations are still negative and often strongly negative at virtually all leads

Table 1
Cross-correlations of prices and output with the same transformation applied to both series

Lag	Canada	France	Germany	Italy	Japan	U.K.	U.S.
Correlations of detrended prices and output							
4	− 0.43	0.24	0.53	− 0.55	0.53	− 0.63	− 0.61
3	− 0.47	0.19	0.49	− 0.60	0.51	− 0.67	− 0.64
2	− 0.50	0.14	0.44	− 0.64	0.49	− 0.71	− 0.67
1	− 0.53	0.08	0.38	− 0.69	0.48	− 0.74	− 0.69
0	− 0.57	0.02	0.34	− 0.73	0.46	− 0.77	− 0.71
− 1	− 0.59	− 0.04	0.28	− 0.74	0.41	− 0.79	− 0.71
− 2	− 0.60	− 0.09	0.23	− 0.75	0.37	− 0.80	− 0.71
− 3	− 0.62	− 0.14	0.18	− 0.75	0.33	− 0.81	− 0.71
− 4	− 0.63	− 0.18	0.14	− 0.75	0.29	− 0.81	− 0.71
Std. err.	0.08	0.11	0.09	0.09	0.08	0.08	0.08
Correlations of detrended prices and output (segmented trend for output)							
4	− 0.13	− 0.07	0.67	− 0.45	0.82	− 0.25	− 0.14
3	− 0.16	− 0.13	0.63	− 0.50	0.83	− 0.30	− 0.16
2	− 0.19	− 0.18	0.59	− 0.55	0.83	− 0.34	− 0.19
1	− 0.23	− 0.23	0.55	− 0.60	0.83	− 0.39	− 0.22
0	− 0.27	− 0.28	0.51	− 0.65	0.83	− 0.44	− 0.25
− 1	− 0.30	− 0.31	0.46	− 0.67	0.79	− 0.47	− 0.27
− 2	− 0.32	− 0.33	0.41	− 0.68	0.75	− 0.51	− 0.28
− 3	− 0.35	− 0.34	0.36	− 0.69	0.70	− 0.54	− 0.29
− 4	− 0.37	− 0.34	0.32	− 0.70	0.65	− 0.56	− 0.30
Std. err.	0.08	0.11	0.09	0.09	0.08	0.08	0.08
Correlations of HP-filtered prices and output							
4	0.23	− 0.23	0.40	0.12	0.16	− 0.07	− 0.00
3	0.17	− 0.35	0.37	0.01	0.03	− 0.20	− 0.01
2	0.08	− 0.45	0.26	− 0.14	− 0.05	− 0.35	− 0.02
1	− 0.02	− 0.59	0.08	− 0.31	− 0.16	− 0.45	− 0.06
0	− 0.17	− 0.69	− 0.06	− 0.47	− 0.22	− 0.53	− 0.14
− 1	− 0.20	− 0.68	− 0.20	− 0.52	− 0.31	− 0.51	− 0.25
− 2	− 0.26	− 0.65	− 0.26	− 0.56	− 0.37	− 0.45	− 0.38
− 3	− 0.27	− 0.52	− 0.28	− 0.62	− 0.38	− 0.35	− 0.49
− 4	− 0.28	− 0.37	− 0.24	− 0.61	− 0.38	− 0.25	− 0.56
Std. err.	0.08	0.11	0.09	0.09	0.08	0.08	0.08

Table 1 (continued)

Correlations of first differences of prices and output

4	0.03	0.05	0.11	− 0.04	0.22	− 0.11	− 0.09
3	− 0.01	− 0.10	0.14	− 0.04	− 0.01	− 0.04	− 0.10
2	− 0.01	0.01	0.10	− 0.06	0.10	− 0.15	− 0.07
1	0.06	− 0.16	− 0.04	− 0.13	− 0.06	− 0.12	− 0.01
0	− 0.26	− 0.29	0.02	− 0.30	0.09	− 0.23	− 0.03
− 1	− 0.07	− 0.15	− 0.11	− 0.14	− 0.02	− 0.13	− 0.09
− 2	− 0.09	− 0.29	− 0.07	− 0.09	− 0.06	− 0.16	− 0.16
− 3	− 0.10	− 0.17	− 0.12	− 0.25	− 0.01	− 0.09	− 0.23
− 4	− 0.10	− 0.10	0.09	− 0.22	− 0.03	− 0.12	− 0.25
Std. err.	0.08	0.11	0.09	0.09	0.08	0.08	0.08

Lag 4 indicates a correlation of the filtered prices series with the fourth lag of the filtered output series. A negative lag denotes a lead. The sample periods are as follows: Canada 47:1–91:4, France 70:1–91:4, Germany 60:1–91:4, Italy 60:1–91:4, Japan 55:2–91:4, U.K. 55:1–91:4, U.S. 47:1–91:4. For the segmented trend, a break in the trend was allowed in 1974:1 for each country. The standard errors are approximate standard errors computed under the null that the true correlation coefficient is zero.

and lags.[7] Another feature of this panel is that the contemporaneous correlations between the HP-filtered price and output series are negative for all countries and, except in the case of Germany, strongly so.

Finally, the fourth panel of Table 1 reports correlations between the first differences of prices and output. These results are interpretable as the correlations between inflation and the rate of growth of output. An important distinction needs to be made here between these correlations and the correlations between inflation and the cyclical component of output that we present later. Although first differencing the level of output does yield a stationary series for the countries in our sample, this stationary series does not generally correspond to a measure of the cyclical component of output.[8]

[7]For the U.S., these negative correlations are much larger in absolute magnitude in the post-Korean war period of 1954: 1–1989:4 examined by Kydland and Prescott (1990). This is consistent with the findings of Wolf (1991), that the price level in the U.S. has been more strongly countercyclical after the early 1970s. The contemporaneous correlation between HP-filtered prices and output is − 0.49 over the period 1954:1–1991:4; Kydland and Prescott (1990) report a correlation of − 0.55 over the period 1954: 1–1989:4.

[8]This holds regardless of whether output has a stochastic or a deterministic trend. Consider the case where output has an autoregressive stationary component and a permanent component that can be modelled as a random walk. Taking the first difference of output will then yield the sum of the innovation in the permanent component and the first difference of the stationary component. Hence, the first difference of output may turn negative even when output is above its trend level.

The results in this bottom panel are more mixed. A few correlations for Canada, France, Germany, and Japan are positive but are mostly small in magnitude. Most of the correlations for these countries and all of the correlations for Italy, the U.K., and the U.S. are still negative and often significantly negative.

This table is generally supportive of the hypothesis of countercyclical price behavior for the G-7 in the postwar period, although the detrending technique does appear to influence the nature of the results for some countries. We interpret these results as confirming for postwar quarterly data the countercyclical price behavior that Backus and Kehoe (1992) and Smith (1992) find for industrialized countries using annual data and a longer sample period. The results also provide strong confirmation for the countercyclical price behavior in the U.S. reported by Kydland and Prescott (1990) and Cooley and Ohanian (1991).

The authors cited above interpret the countercyclical behavior of the price level as supportive of supply-determined models of the cycle. However, as discussed above, many conventional business cycle models, including most demand-driven models of the cycle, imply that it is the inflation rate or the change in (the logarithm of) the price level that is procyclical rather than the price level itself. To investigate if the countercyclical behavior of the aggregate price level in the G-7 carries over to countercyclical inflation, we now turn to an examination of the correlations between inflation and cyclical output.

2.2. *Correlations between inflation and cyclical output*

As discussed above, output and the price level in the G-7 are clearly non-stationary series. For inflation, the evidence is far less conclusive. In the Appendix, we examine the stationarity of the inflation rate in our sample. The results there lead us to conclude that inflation is stationary in all G-7 countries over our sample period. Hence, in examining the comovement of inflation with the cycle, we report correlations between the levels of inflation and various measures of the cyclical component of output.

In addition to the detrending techniques used earlier to obtain the cyclical component of output, we use one additional method developed by Blanchard and Quah (1989). Under the maintained assumption that output has a unit root, the Blanchard–Quah decomposition is a method of decomposing output into its permanent and temporary components without restricting the permanent component of output to be a random walk.[9] In principle, any variable that is

[9]Blanchard and Quah (1989) interpret their technique as providing a decomposition of output fluctuations into demand and supply shocks. We, however, view their methodology as simply providing a decomposition of output into its cyclical and permanent components with a less restrictive assumption on the permanent component than the Beveridge–Nelson (1981) decomposition (which models the permanent component as a random walk).

stationary and is affected by the same shocks as output could be used in the Blanchard–Quah decomposition. Given the focus of this paper, we decompose output fluctuations using a bivariate vector autoregression with the inflation rate and the first difference of output.

The first panel of Table 2 reports the correlations between inflation and output detrended using a linear trend. Virtually all of the correlations in this

Table 2
Correlations of inflation and various measures of cyclical output.

Lag	Canada	France	Germany	Italy	Japan	U.K.	U.S.
Correlations of inflation and detrended output							
4	0.38	0.57	0.25	0.42	0.38	0.53	0.36
3	0.37	0.57	0.27	0.42	0.39	0.51	0.33
2	0.37	0.58	0.32	0.42	0.39	0.45	0.29
1	0.38	0.57	0.27	0.40	0.40	0.41	0.28
0	0.32	0.52	0.30	0.36	0.40	0.32	0.27
− 1	0.29	0.48	0.28	0.33	0.39	0.26	0.25
− 2	0.25	0.41	0.26	0.31	0.38	0.19	0.20
− 3	0.22	0.37	0.24	0.25	0.37	0.17	0.15
− 4	0.19	0.34	0.27	0.20	0.36	0.12	0.09
Std. err.	0.08	0.11	0.09	0.09	0.08	0.08	0.08
Correlations of inflation and detrended output (segmented trend)							
4	0.39	0.46	0.19	0.44	0.19	0.60	0.37
3	0.40	0.42	0.22	0.44	0.22	0.63	0.37
2	0.41	0.37	0.29	0.45	0.25	0.60	0.38
1	0.45	0.31	0.25	0.44	0.32	0.59	0.39
0	0.41	0.25	0.28	0.40	0.37	0.48	0.40
− 1	0.39	0.17	0.27	0.37	0.42	0.44	0.39
− 2	0.37	0.08	0.26	0.35	0.44	0.38	0.36
− 3	0.35	0.05	0.23	0.31	0.46	0.33	0.31
− 4	0.32	0.02	0.27	0.26	0.48	0.27	0.25
Std. err.	0.08	0.11	0.09	0.09	0.08	0.08	0.08
Correlations of inflation and HP-filtered output							
4	0.13	0.21	0.08	0.19	0.25	0.25	0.05
3	0.17	0.18	0.16	0.23	0.18	0.27	0.06
2	0.19	0.22	0.23	0.25	0.22	0.19	0.09
1	0.27	0.18	0.17	0.23	0.12	0.15	0.15
0	0.13	0.05	0.19	0.09	0.17	0.02	0.19
− 1	0.09	0.02	0.11	0.06	0.11	− 0.04	0.19
− 2	0.04	− 0.12	0.06	0.07	0.03	− 0.12	0.16
− 3	− 0.01	− 0.16	− 0.01	− 0.03	0.00	− 0.11	0.09
− 4	− 0.06	− 0.15	0.07	− 0.12	− 0.04	− 0.16	0.02
Std. err.	0.08	0.11	0.09	0.09	0.08	0.08	0.08

Table 2 (continued)

Correlations of inflation and cyclical output obtained using a bivariate Blanchard–Quah decomposition with output and inflation

4	0.44	0.49	0.49	0.25	0.29	0.16	0.49
3	0.49	0.49	0.31	0.30	0.21	0.59	0.51
2	0.56	0.51	0.32	0.34	0.34	0.61	0.54
1	0.63	0.49	0.23	0.31	0.43	0.57	0.58
0	0.74	0.42	0.49	0.40	0.66	0.79	0.71
−1	0.72	0.39	0.46	0.38	0.63	0.63	0.71
−2	0.64	0.34	0.35	0.35	0.60	0.57	0.68
−3	0.56	0.33	0.33	0.25	0.56	0.53	0.60
−4	0.45	0.32	0.32	0.16	0.51	0.47	0.53
Std. err.	0.08	0.11	0.09	0.09	0.08	0.08	0.08

Lag 4 indicates the correlation between inflation and the fourth lag of cyclical output. A negative lag denotes a lead. Also see notes to Table 1.

panel are strongly positive indicating that inflation is procyclical. However, as discussed above, simple linear detrending of output may yield spurious results. The second panel of this table reports the correlations between inflation and output detrended using a segmented linear trend. Again, the correlations at most leads and lags, including the contemporaneous correlations, are strongly positive for all of the G-7 countries. For all countries except Japan and the U.S., the correlations between cyclical output and inflation peak at a positive lag, suggesting a lagged effect of cyclical output on inflation. For Japan, the peak positive correlation occurs at lead 4 while, for the U.S., it occurs at lag 0.

The results in the third panel, using HP-filtered output, are similar to the second panel except that the contemporaneous correlations are now strongly positive only for Germany, Japan, and the U.S. A few negative correlations appear at the leads in this panel, but most of these are small in magnitude. Most of the correlations between inflation and HP-filtered output are positive and often significantly positive. The peak positive correlations occur at the lags for all countries except the U.S., for which the contemporaneous correlation is the largest. The largest positive correlations occur at lag 0 for the U.S., the first lag for Canada, the second lag for France, Germany, and Italy, the third lag for the U.K., and the fourth lag for Japan.

Finally, the correlations between inflation and cyclical output obtained from a Blanchard–Quah decomposition are presented in the bottom panel of Table 2. The correlations in this panel are all resoundingly positive.

The results in Table 2 corroborate the conventional wisdom of procyclical variation in inflation. In the postwar sample that we have examined for the G-7 countries, the countercyclicality of the price level clearly does not translate into countercyclical inflation.[10] In particular, for the U.S., it is striking that every single correlation between the cyclical components of prices and output is negative, while all the correlations between inflation and the cyclical component of output are positive.[11]

Fig. 1 plots inflation (annualized rates) and the HP-filtered price level for the U.S. This is a striking picture. It clearly shows that the two variables exhibit very different behavior and often move in opposite directions. Fig. 2 contains time series plots of inflation (annualized rates) and two measures of cyclical output – output detrended using a segmented linear trend and using the HP filter. The positive correlation between inflation and cyclical output in the U.S. is apparent in this figure. Fig. 3 plots HP-filtered prices and the same measures of cyclical output as in Fig. 2. A closer inspection of Figs. 2 and 3 also hints at another important detail – that the relation between prices and inflation and the business cycle may not have been stable over time in the postwar U.S.[12] For instance, Wolf (1991) has shown that, for the postwar U.S., the price level has been more strongly countercyclical after the early 1970s. The behavior of inflation over the cycle appears to be much less interval- dependent. Over various subperiods that we have examined for the U.S., inflation–output correlations are consistently positive although varying in magnitude. We do not pursue this feature of the data in this paper but it suggests the need for further caution in interpreting stylized facts regarding the cyclical behavior of prices.

Although we have examined correlations of inflation and the price level with alternative measures of cyclical output, it is also useful to examine the robustness of our results using alternative indicators of the cycle. Accordingly, we now turn to an examination of correlations using the unemployment rate as an indicator of the cycle.

[10]In line with our results for the G-7, Hassler et al. (1992) report that, in postwar annual data for Sweden, the cyclical components of prices and output have a contemporaneous negative correlation, while the correlation between inflation and cyclical output is positive.

[11]In related work (Chadha and Prasad, 1993), we have also used other stochastic detrending techniques such as the Beveridge–Nelson (1981) decomposition for U.S. output. In that paper, we report strong positive correlations between U.S. inflation and cyclical output, when the latter variables is obtained using (i) a Beveridge–Nelson decomposition from an ARIMA (5, 1, 1) process fitted to the level of output and also (ii) a bivariate Blanchard–Quah decomposition using the first difference of output and the unemployment (corresponding to what Blanchard and Quah refer to as the 'base case' in their paper).

[12]Using annual data for ten industrialized countries including the U.S., Backus and Kehoe (1991) have documented substantial differences among the prewar, interwar, and postwar correlations between HP-filtered prices and output.

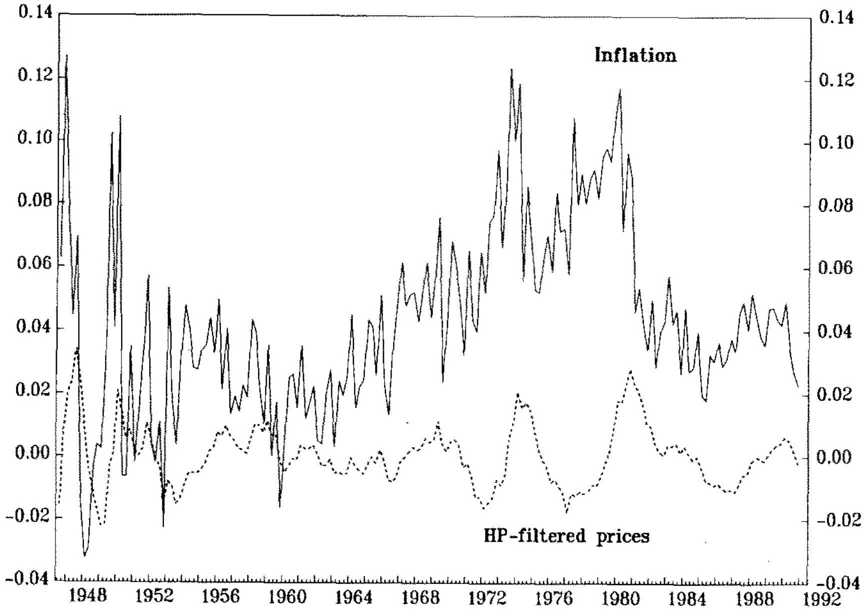

Fig. 1. Inflation and HP-filtered prices in the U.S., 1947:2–1991:4.

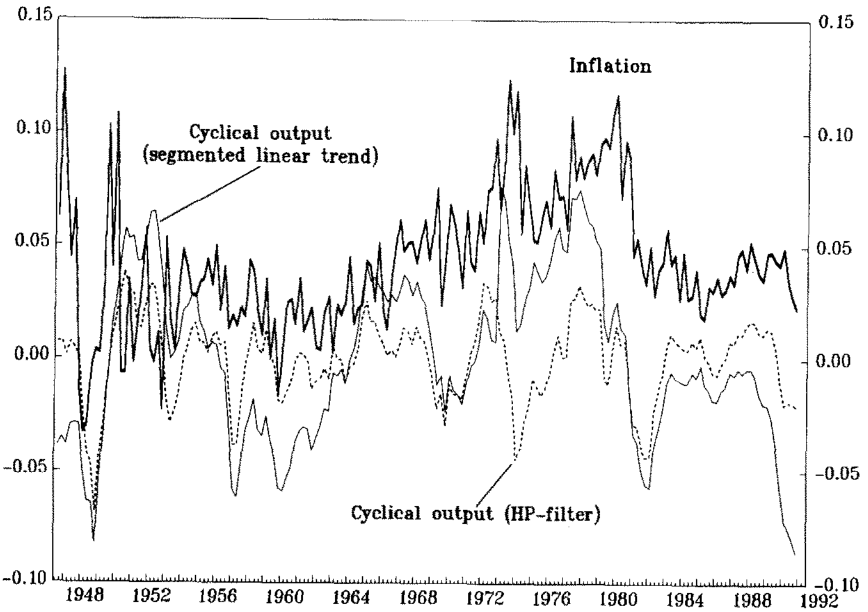

Fig. 2. Inflation and cyclical output in the U.S., 1947:2–1991:4.

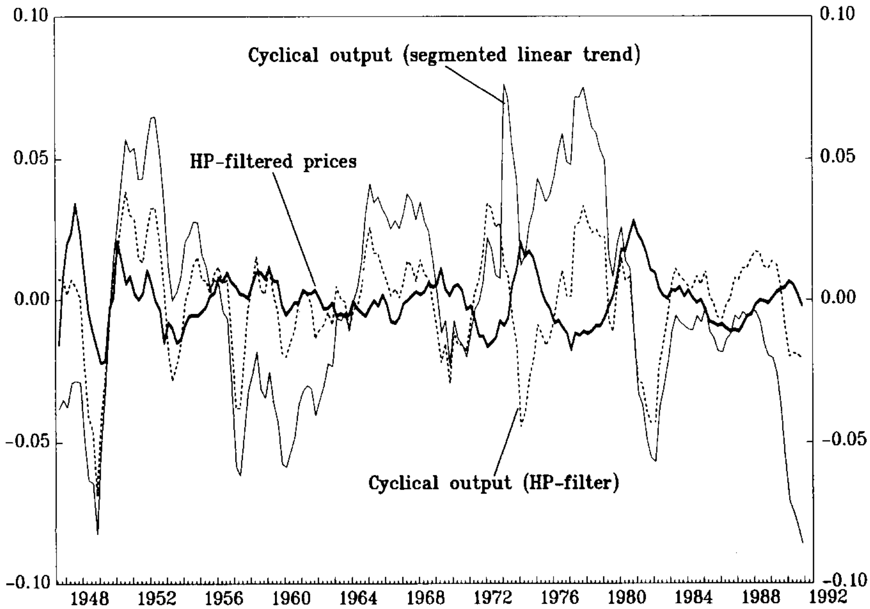

Fig. 3. HP-filtered prices and cyclical output in the U.S., 1947:2–1991:4.

3. Price and inflation correlations with unemployment

The unemployment rate provides an alternative measure of the cycle and is generally strongly negatively correlated with cyclical output. However, it is sometimes the case that the goods and labor markets provide different indications about the stage of the business cycle that an economy is in. Hence, it is useful to examine whether the correlations reported above are robust to changes in the cyclical indicator. Further, in the context of the Phillips curve literature, the relation between inflation and unemployment is of considerable interest.

A number of studies have concluded that the unemployment rate is non-stationary in most European countries including those represented in our sample (e.g., Blanchard and Summers, 1987). Rather than use a variety of detrending procedures for the unemployment rate, we use only the HP filter to detrend the unemployment rate in all countries. Note that, since unemployment is negatively correlated with the cyclical component of output, a positive correlation between unemployment and any other variable implies that that variable moves countercyclically, while a negative correlation indicates procyclical variation. The seasonally adjusted quarterly unemployment

Table 3
Correlations of prices and inflation with cyclical unemployment.

Lag	Canada	France	Germany	Italy	Japan	U.K.	U.S.
Correlations of HP-filtered prices and unemployment							
4	− 0.42	− 0.12	− 0.15	− 0.30	− 0.28	− 0.37	− 0.03
3	− 0.37	− 0.05	− 0.07	− 0.31	− 0.27	− 0.26	− 0.06
2	− 0.29	0.09	0.01	− 0.30	− 0.24	− 0.12	− 0.09
1	− 0.18	0.25	0.07	− 0.25	− 0.14	0.04	− 0.07
0	− 0.04	0.38	0.15	− 0.18	− 0.02	0.20	0.00
− 1	0.12	0.47	0.27	− 0.04	0.13	0.34	0.14
− 2	0.28	0.51	0.33	0.07	0.29	0.44	0.30
− 3	0.41	0.50	0.37	0.19	0.44	0.48	0.45
− 4	0.51	0.44	0.37	0.31	0.54	0.48	0.56
Std. err.	0.09	0.12	0.10	0.09	0.09	0.09	0.08
Correlations of inflation and HP-filtered unemployment rate							
4	− 0.15	− 0.12	− 0.16	− 0.04	− 0.05	− 0.22	0.01
3	− 0.21	− 0.18	− 0.15	− 0.07	− 0.08	− 0.26	− 0.00
2	− 0.27	− 0.21	− 0.13	− 0.11	− 0.18	− 0.29	− 0.06
1	− 0.33	− 0.18	− 0.13	− 0.15	− 0.22	− 0.30	− 0.12
0	− 0.37	− 0.16	− 0.17	− 0.23	− 0.24	− 0.28	− 0.20
− 1	− 0.35	− 0.08	− 0.11	− 0.19	− 0.28	− 0.21	− 0.23
− 2	− 0.31	0.01	− 0.07	− 0.20	− 0.25	− 0.12	− 0.22
− 3	− 0.25	0.06	− 0.03	− 0.22	− 0.17	− 0.07	− 0.15
− 4	− 0.18	0.11	0.01	− 0.16	− 0.02	0.02	− 0.06
Std. err.	0.09	0.12	0.10	0.09	0.09	0.09	0.08

Lag 4 indicates the correlation between HP-filtered prices or the inflation rate and the fourth lag of the HP-filtered unemployment rate. A negative lag denotes a lead. Seasonally adjusted quarterly unemployment rate data were available as follows: Canada 57:1–91:4, France 72:1–91:4, Germany 69:1–91:4, Italy 63:1–91:4, Japan 60:1–91:4, U.K. 57:1–91:4, U.S. 48:1–91:4. The standard errors reported here are approximate standard errors computed under the null hypothesis that the true correlation coefficient is zero.

rate data for the G-7 used in this section were obtained from the IMF's CEI databank.[13]

The first panel of Table 3 presents correlations between HP-filtered prices and the HP-filtered unemployment rate. The results are mixed. The contemporaneous correlations between HP-filtered prices and unemployment are positive

[13]This dataset will be made available upon request. We chose this source since our only available alternative for quarterly unemployment data, the OECD databank, had data starting only in 1970. For the period over which they overlap, the CEI and OECD databanks are consistent with each other.

for France, Germany, and the U.K, confirming the countercyclical behavior of prices. They are close to zero for the U.S. and negative for Canada, Italy, and Japan. For Canada, Italy, Japan, and the U.K., a number of significant negative correlations appear at the lags. However, for all countries in the sample, the correlations are strongly positive at the leads. These results fail to provide a strong confirmation that the price level is countercyclical in the G-7, although many of the correlations do support this hypothesis.

The lower panel of Table 3 presents correlations between inflation and the HP-filtered unemployment rate. The correlations are strongly negative for virtually all countries and at most leads and lags. Further, the contemporaneous correlation between cyclical unemployment and the inflation rate is now significantly negative for most of the countries. The results in this panel bolster the results from Table 2 that the inflation rate in the G-7 countries has been procyclical in the postwar period.

It is worth noting a feature of these results. For most countries, the correlations of inflation with cyclical unemployment peak at or very close to lag 0. In the case of the correlations with cyclical output, however, the positive correlations generally peaked at higher lags. For instance, as noted before, the largest positive correlation is between inflation and the second lag of HP-filtered output for Germany and Italy. For both these countries, the largest negative correlation between inflation and HP-filtered unemployment occurs at lag 0. This is consistent with the view that employment growth typically lags the cycle, as measured by output deviations from trend, in most industrialized countries. Consequently, the unemployment rate tends to lag the business cycle in a similar fashion, resulting in the peak correlations between inflation and cyclical unemployment occurring at smaller lags than the peak correlations between inflation and cyclical output.[14]

Another feature of these results is that the largest positive correlations between inflation and HP-filtered output are numerically of the same order of magnitude as the largest correlations between inflation and HP-filtered unemployment for most countries, although these two sets of correlations peak at different lags. For instance, for the U.S., the strongest correlation between inflation and HP-filtered output is $+0.19$ (at lag 0) while the strongest negative correlation between inflation and cyclical unemployment is -0.23 (at lead 1).

[14]However, the relationship between employment growth and changes in the unemployment rate may not always be tight due to cyclical changes in the labor force participation rate.

4. Concluding remarks

The main finding of this paper is that, in the G-7, inflation is positively correlated with various measures of the cyclical component of output. We also confirmed, using postwar quarterly data for the G-7, the countercyclical behavior of the price level reported recently by a number of authors. Using unemployment as an alternative indicator of the business cycle, we reported corroborating evidence for the procyclical behavior of inflation but were not able to obtain strong confirmation of countercyclical price variation.

Our results suggest that the cyclical behavior of the price level is consistent with the predictions of supply-determined models of the business cycle, including real business cycle models. However, the procyclicality of the inflation rate rather than the price level is a key implication of many other classes of business cycle models, including conventional demand-determined models of the cycle. The procyclical behavior of inflation that we have found is consistent with these models. Thus, our findings suggest that the cyclical behavior of the price level and inflation do not provide conclusive grounds for rejecting either demand-determined or supply-determined models of the cycle.

Further research is necessary to reconcile the findings reported in our paper. This paper has shown that, in the process of analyzing the implications of alternative business cycle models, it is essential to differentiate clearly between a model's predictions for inflation and for the cyclical component of prices.

Appendix: Stationarity of the rate of inflation

This appendix examines the stationarity of the inflation rate in postwar quarterly data for the G-7. Table A.1 reports the results of standard Dickey–Fuller (DF) and Augmented Dickey–Fuller (ADF) regressions for the inflation rate. All of the regressions included a constant. The table also reports the significance level of the Q-statistic for the estimated residuals from each regression. The Q-statistic provides a test of the hypothesis that the residuals from the regression are serially uncorrelated.

The absolute t-statistics from the DF regressions are all large enough to suggest that the hypothesis that there is a unit root in the inflation rate for the G-7 can be rejected at the 5 percent level. The ADF regressions with one lag confirm this result except for France where the ADF statistic drops to -2.38. When the ADF regressions are extended to allow for four lags, the t-statistics are higher than the 5 percent critical value for all countries except France, Germany, and the U.K. For the latter two countries, the t-statistic is just below the 5 percent critical value and the null hypothesis of a unit root can be rejected at the 10 percent level. However, for France, the t-statistic drops to -1.15, well

Table A.1
Dickey Fuller (DF) and Augmented Dickey–Fuller (ADF) tests for a unit root in inflation

	DF statistic	Q statistic signif. level	AD statistic 1 lag	Q statistic signif. level	ADF statistic 4 lags	Q statistic signif. level
Canada	− 6.81	0.00	− 4.68	0.00	− 3.86	0.39
France	− 4.15	0.00	− 2.38	0.01	− 1.15	0.68
Germany	− 14.22	0.00	− 7.30	0.00	− 2.84	0.57
Italy	− 8.84	0.00	− 5.04	0.02	− 2.90	0.16
Japan	− 6.86	0.00	− 4.61	0.00	− 3.05	0.08
U.K.	− 6.24	0.00	− 3.43	0.70	− 2.81	0.80
U.S.	− 5.92	0.02	− 3.87	0.63	− 3.34	0.49

The regressions include a constant. The critical value at the 5 percent significance level is −2.89 from Fuller (1976). The Q-statistic tests whether the regression residuals are white noise. A significance level higher than 0.10 for the Q-statistic indicates that the hypothesis that the residuals are white noise cannot be rejected at the 10 percent level of significance.

Table A.2
Estimates of 'normalized spectral density' at frequency zero for inflation

Number of lags	Canada	France	Germany	Italy	Japan	U.K.	U.S.
6	0.30	0.20	0.11	0.17	0.23	0.23	0.27
12	0.19	0.13	0.07	0.08	0.12	0.14	0.15
24	0.13	0.12	0.05	0.05	0.08	0.09	0.10
48	0.12	0.08	0.03	0.03	0.05	0.06	0.08
96	0.09		0.02	0.02	0.02	0.04	0.05
168	0.04						0.02

The first column refers to the number of lagged autocorrelations used to construct the estimate of the 'normalized spectral density' at frequency zero. The maximum number of lagged autocorrelations used in this construction varies across countries due to differences in the length of the sample periods.

below even the 10 percent critical value. For each of the ADF regressions with 4 lags, the Q-statistic has a significance level greater than 5 percent. This indicates that the hypothesis that the residuals from these regressions are white noise cannot be rejected at the 5 percent level of significance, suggesting that allowing for a maximum lag length of four in the ADF regressions is sufficient.

As a complement to the above tests, we also employed a stationarity test that is based on estimating the ratio of 2Π times the spectral density of the first difference of a series at frequency zero, to the variance of its first difference (see Huizinga, 1987, for a similar application of these tests to exchange rates). Note that this ratio, referred to below as the 'normalized density', is constructed using estimates of the spectral density at frequency zero rather than the spectral density *function*. By examining autocorrelations at long lags of the series, the test has the potential advantage of being able to detect slowly evolving changes in the series rather than relying solely on the point estimates yielded by the DF and ADF tests. To interpret the results, two observations are useful. First, when the level of a series follows any stationary stochastic process, this ratio or normalized density will tend toward zero as the number of lags used in its construction increases (goes to infinity). Second, for any series that is integrated of order one, the ratio should converge to the ratio of the variance of changes in the permanent component to the variance of total changes in the variable.

Table A.2 reports estimates of the normalized density for the inflation rate in the G-7. For all countries, the estimated normalized density falls below unity and declines as the number of lagged autocorrelations employed increases and approaches zero at the maximum number of lags employed. For instance, in the case of Canada, the estimate drops from 0.30 using six lagged autocorrelations to 0.04 using 168 lagged autocorrelations. In the case of France, although the ADF regressions did not lead to the rejection of the hypothesis of a unit root in inflation, the results in this table suggest that inflation may in fact be mean-reverting. As is well-known, the low power of the DF and ADF tests makes it difficult to reject the null hypothesis of a unit root if a large autoregressive root is present.

These results support the hypothesis that the inflation rate is stationary in postwar quarterly data for all of the G-7 countries. While the fact that the normalized density approaches zero does not establish that the series are stationary, these results suggest that, even if there exist permanent stochastic components in the rate of inflation for the G-7, these components are relatively small.

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