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# Interpreting the Cyclical Behavior of Prices

BANKIM CHADHA and ESWAR PRASAD*

*This paper argues that determining the cyclical behavior of prices by applying the same stationarity-inducing transformation to the levels of both output and prices, and examining the correlations of the resulting series, can be misleading. A more appropriate procedure is to examine the correlations between the rate of inflation and the level of the cyclical component of output. In postwar U.S. data, the correlations between similarly transformed price and output data are consistently and often strongly negative. The rate of inflation, however, is consistently and usually strongly positively correlated with various measures of the cyclical component of output.* [JEL E31, E32]

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. Closely related to this question is the issue of the co-movement of prices with the cyclical component of output. If temporary movements of output result primarily from shocks to demand, prices would be expected to be procyclical; if the movements result from shocks to supply, prices would be expected to be countercyclical. It is widely perceived that temporary movements in output are

* Bankim Chadha is an Economist in the IMF's Research Department. He received his doctorate from Columbia University.
Eswar Prasad is an Economist in the Western Hemisphere Department. He received his doctorate from the University of Chicago.
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associated with shocks to demand, while longer-term movements are associated with movements in supply. Blanchard and Quah (1989, p. 656), for example, in decomposing fluctuations in aggregate output into temporary and permanent components "interpret the disturbances that have a temporary effect on output as being mostly demand disturbances, and those that have a permanent effect on output as mostly supply disturbances." It has also been widely perceived that price movements are procyclical.[1] Lucas (1976, p. 19), for example, 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."

Recent work by Kydland and Prescott (1990), Backus and Kehoe (1992), Cooley and Ohanian (1991), and Smith (1992) has reexamined the cyclical variation of prices. These authors provide evidence that contradicts the conventional wisdom that price movements are procyclical. In the most comprehensive of these studies 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.[2] The cross-correlations of the resulting series suggest 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 and suggests that even temporary movements in output may be the result of supply disturbances.

Aggregate output and the price level in the United States are clearly nonstationary series. Examining the co-movements of any two nonstationary series by the raw correlations of their levels will yield spurious results. In the absence of theoretical priors that one is willing to impose on the data, it seems innocuous, if not logical, to apply the same transformation to both series in order to render them stationary and then examine the co-movements of the resulting series. This is in fact standard practice, with the dominant procedures for transforming the data being those employed by Cooley and Ohanian (1991).

This paper argues that to examine the variation of prices with cyclical

[1] This positive correlation is often associated with some form of the "Phillips curve." We do not pursue this interpretation in this paper, although the findings are clearly relevant.

[2] Hodrick and Prescott (1980). All series are first transformed into logarithms.

movements in output, the three procedures mentioned could yield misleading conclusions. Our argument focuses on the response of prices to cyclical movements in output when the cyclical movements in output are associated with movements in demand. Consider an economy where output is demand determined in the short run and prices are sticky in nominal terms.[3] In such an economy any (nominal or real) shock that increases aggregate demand will, on impact, raise output above its long-run trend level.[4,5] Over time, excess demand causes the price level to rise, which reduces aggregate demand and gradually returns output to the level dictated by its long-run trend. In this economy, where prices are completely procyclical by construction, there will be a positive correlation between *changes* in the price level—the rate of inflation—and the *level* of the cyclical component of output, that is the deviation of observed output from a longer-term (deterministic or stochastic) trend. We show that alternative measures of the co-movement of prices with the cyclical component of output—such as the three standard procedures employed by Cooley and Ohanian (1991)—can be misleading when output is demand determined in the short run.

If short-run movements in output were supply determined, however, some of the standard procedures could be valid. Consider an economy where the cyclical component of output results from exogenous temporary shocks to supply and where prices are flexible. In response to a positive innovation in the cyclical component of the supply of output, the price level will fall. Subsequently, as the shock dissipates over time, the price level will rise. There will, therefore, be a negative correlation between the cyclical components of the price level and output. If long-run movements in output were described by a time trend, the cyclical components of prices and output could be obtained by detrending the two series.

The two alternative hypotheses of the short-run determination of output predict correlations between alternative transformations of the data. Under the assumption that output is demand determined, one would

[3] For output to be demand determined and to differ nontrivially from aggregate supply, there must exist some friction, imperfect information, or other coordination failure in the economy that causes prices to deviate from their "flexible" levels, which, by definition, equilibrate demand and supply. One form of such a friction that is often appealed to is the existence of sticky nominal wages or prices. A celebrated example is that of rational staggered nominal wage contracting developed by John Taylor (1979, 1980).

[4] This "trend" need not be deterministic.

[5] As a standard example of such a shock consider the effect of an unanticipated increase in the level of the nominal money stock. The traditional channel by which output expands is that with sticky prices the nominal increase in the money supply translates into a real increase, placing downward pressure on interest rates and raising aggregate demand.

expect to find a positive correlation between the change in prices (the rate of inflation) and cyclical movements in output. Under the assumption that output is supply determined, on the other hand, one would expect to find a negative correlation between cyclical movements in the levels of prices and output. The supply-determined output model, however, also implies a strong prediction for the correlation between the rate of inflation and cyclical movements in output. On impact of a positive supply shock, as the price level falls, so will the rate of inflation, implying a negative contemporaneous correlation. Subsequently, however, since the price level rises over time as the shock dissipates, the inflation rate will jump up and then gradually fall to its long-run value. While the correlation between inflation and the cyclical component of output will, therefore, depend on the particular lead or lag examined, a key feature of the cross-correlations is that there will be at least some significant negative correlations implied by the response of the price level to an innovation in the cyclical component of supply. Examining the cross-correlations between inflation and the cyclical component of output should therefore yield a better test of the cyclical behavior of prices than the standard procedures, which could be misleading if output is demand determined.

In the following section, we examine the co-movement of prices and output for the postwar period in the United States. We find that correlations are negative at various lags and leads when the first differences of the series, their deviations from a deterministic trend, or their filtered values using the Hodrick-Prescott procedure are compared. This is consistent with the evidence presented by Kydland and Prescott (1990), Backus and Kehoe (1992), Smith (1992), and particularly Cooley and Ohanian (1991), which suggests that prices are countercyclical in the postwar United States. We then examine the cross-correlations between the rate of inflation and the cyclical component of output under alternative hypotheses about the nature of long-run movements in output. Under the assumption that long-run movements in output can either be described as deterministic functions of time or be picked up by the Hodrick-Prescott filter and the cyclical component of output is measured by either detrended or Hodrick-Prescott filtered output, the correlations are consistently and often significantly positive suggesting that prices are indeed procyclical. Under the assumption that output has a stochastic trend, we measure the cyclical component of output using both the Beveridge-Nelson (1981) and Blanchard-Quah (1989) decompositions. The correlations are still typically positive and often significantly so.

We then illustrate the effects of applying alternative transformations to prices and output by examining two simple stylized macroeconomic models: a sticky-price demand-driven model, and a supply-driven model with flexible prices. Stochastic simulations of the models are performed, and the cross-correlations between prices and output in the generated data are examined with alternative transformations of the data. The simulations show that when output is demand driven the correlation between inflation and the cyclical component of output reflects most accurately the procyclical behavior of prices. Alternative correlations, such as those between detrended and first-differenced data, are shown often to yield negative cross-correlations. Simulations of the supply-driven model show that while correlations between detrended data accurately reflect the countercyclical behavior of prices the model also yields various negative correlations between inflation and the cyclical component of output. The final section contains concluding remarks.

## I. Output and Prices in the Postwar United States

This section provides evidence on the nature of the co-movement of prices with the cyclical component of output in the postwar United States at a quarterly frequency. Cross-correlations of various transformations of prices and output are compared in an attempt to ascertain the cyclical behavior of prices.

Aggregate output is clearly a nonstationary series. Characterization of the form of the nonstationarity is important, since this determines the appropriate measure of the cyclical component of output. There are two possibilities. The first possibility is that output is stationary around a deterministic time trend. Under this assumption, an estimate of the long-run component of output can be obtained as the series of fitted values from a regression of output on a constant and a time trend; the residual series then provides an estimate of the cyclical component of output. The second possibility is that output has a unit root or is characterized by stochastic nonstationarity. Since the seminal work of Nelson and Plosser (1982), the view that most economic time series are characterized by stochastic rather than deterministic nonstationarity has become increasingly prevalent. Under the assumption that output has a unit root, one method of identifying long-run movements in output is associated with Beveridge and Nelson (1981). They show that any variable that has a unit root can be represented as the sum of a permanent component and a temporary component, where the permanent component is a random walk in which all changes are by definition permanent and the temporary component is some stationary process. Another well-known method of decomposing output under the maintained assumption that output has a unit root is associated with Blanchard and Quah (1989). Their methodology differs from that of Beveridge and Nelson (1981) in that it allows for dynamic effects of disturbances that have permanent effects. They do not, therefore, restrict the "permanent" component of output to be a random walk.

A number of studies have concluded that postwar quarterly GNP in the United States is characterized by unit-root nonstationarity (see, among others, Nelson and Plosser (1982) and Campbell and Mankiw (1987)). However, in an important paper, Perron (1989) finds that the unit-root hypothesis can be rejected in favor of the hypothesis that output is stationary around a deterministic trend that has one break in its slope in 1973. This break coincides with the first oil shock and the start of the productivity slowdown in the United States. The conflicting evidence on the form of the nonstationarity is important since the two alternative views on the form of the nonstationarity of output imply the use of different measures of the cyclical component of output. As Perron (1989) points out, however, it is worth noting that there is a certain observational equivalence between the two views. The unit-root hypothesis literally implies that shocks that have a permanent effect on output occur in every period. The segmented-trend hypothesis implies that there is a once-and-for-all shock at the time of the break in the trend whose effect persists forever. Under either hypothesis, there was a shock in 1973 whose effect still persists.[6]

Whether output is best described by a unit-root process or as a process that is stationary around a segmented time trend remains an unresolved question. Our aim is, therefore, to adopt an eclectic approach to measuring the cyclical component of output. In this paper, we employ various measures of the cyclical component of output: (i) output detrended by regressing the series on a constant and a time trend; (ii) detrended output with a break in the trend in the first quarter of 1973 as suggested by Perron (1989); (iii) Hodrick-Prescott filtered output, a transformation that is popular in the real business cycle literature;[7] (iv) the stationary compo-

[6] From a practical point of view, the shock in this time period seems to dominate the data in terms of permanent shocks over the postwar period. Also see Perron (1988), Cochrane (1988), Christiano and Eichenbaum (1990), and Banerjee, Lumsdaine, and Stock (1990).

[7] The Hodrick-Prescott filter involves a linear transformation of a time series and is obtained as the solution to the following 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 $\lambda = 1{,}600$ as suggested by Prescott (1986). See King and Rebelo (1989) and Cogley and Nason (1991) for an analysis of the properties of the Hodrick-Prescott filter.

nent of output obtained from Beveridge-Nelson decompositions under alternative assumptions of the ARIMA process describing output; and (v) the component of output associated with disturbances that have a temporary effect on output, using two decompositions along the lines of Blanchard and Quah (1989).

The output series used in this study is quarterly real GNP measured in 1982 dollars. To maintain comparability with other recent studies the price series employed is the implicit GNP deflator. All of the estimates computed for the GNP deflator were also computed for the consumer price index (CPI). The results were broadly similar using the CPI and do not significantly affect any of our conclusions; they are, therefore, not reported here. All three series were obtained from the Data Resources Incorporated data bank. Our sample covers the postwar period from 1947:1 to 1989:4.

Table 1 reports the results of using the standard technique of applying the same transformation to both output and prices and examining the cross-correlations of the resulting series.[8] The correlations between the log difference of prices with up to four lags and leads of the first difference of output (first panel) are small and negative for the most part, except at the third and fourth leads where they are significantly negative.[9] The correlations between detrended prices and output (second panel), where the detrending involved removing a constant and a linear time trend, are all significantly negative. When output is modeled as stationary around a segmented time trend, along the lines suggested by Perron (1989), but allowing for a break in the level and the slope of the trend in 1973:1 (third panel), the correlations are also negative, though smaller than in the previous panel. The fourth panel presents the correlations of Hodrick-Prescott filtered prices and output. The correlations are negative at all leads and lags and strongly so at the leads.

The remarkable feature of Table 1 is that irrespective of the transformation employed, each and every reported cross-correlation is negative. While the strength of the negative relationship depends on the measure used, the results in this table support the general conclusion that the cyclical components of prices and output are negatively correlated at a quarterly frequency. There is absolutely no evidence for procyclical price behavior. In fact, there is strong evidence for countercyclical price behavior.

[8] These results are comparable to those in Table 1 of Cooley and Ohanian (1991) except that the sample period has been extended to include more recent data.

[9] Standard errors were computed under the null hypothesis that the two series are uncorrelated.

Table 1. *Cross-Correlations of Prices and Output with the Same Transformation Applied to Both Series, Quarterly Data, 1947:2 to 1989:4*

| Lag | Cross-correlation | Standard error |
|---|---|---|
| Log differences of prices and output | | |
| 4 | −0.10 | 0.08 |
| 3 | −0.10 | 0.08 |
| 2 | −0.07 | 0.08 |
| 1 | −0.06 | 0.08 |
| 0 | −0.07 | 0.08 |
| −1 | −0.08 | 0.08 |
| −2 | −0.13 | 0.08 |
| −3 | −0.24 | 0.08 |
| −4 | −0.24 | 0.08 |
| Detrended prices and output | | |
| 4 | −0.62 | 0.06 |
| 3 | −0.64 | 0.06 |
| 2 | −0.66 | 0.06 |
| 1 | −0.68 | 0.06 |
| 0 | −0.69 | 0.06 |
| −1 | −0.69 | 0.06 |
| −2 | −0.69 | 0.06 |
| −3 | −0.69 | 0.06 |
| −4 | −0.68 | 0.06 |
| Detrended prices and output (trend break for output in 1973:1) | | |
| 4 | −0.07 | 0.08 |
| 3 | −0.08 | 0.08 |
| 2 | −0.09 | 0.08 |
| 1 | −0.09 | 0.08 |
| 0 | −0.10 | 0.08 |
| −1 | −0.10 | 0.08 |
| −2 | −0.10 | 0.08 |
| −3 | −0.10 | 0.08 |
| −4 | −0.09 | 0.08 |
| Hodrick-Prescott filtered prices and output | | |
| 4 | −0.02 | 0.08 |
| 3 | −0.04 | 0.08 |
| 2 | −0.06 | 0.08 |
| 1 | −0.11 | 0.08 |
| 0 | −0.19 | 0.08 |
| −1 | −0.28 | 0.07 |
| −2 | −0.38 | 0.07 |
| −3 | −0.48 | 0.07 |
| −4 | −0.54 | 0.07 |

Note: Lag 4 indicates a correlation of the transformed price series with the fourth lag of the transformed output series. A negative lag denotes a lead.

The main argument of this paper is that determining the cyclical behavior of prices by examining the correlations between similarly transformed price and output series, as is done in Table 1, is likely to be misleading when output is demand determined in the short run. If output is demand determined, then the appropriate comparison is between the rate of inflation and the cyclical component of output, and one would expect to find positive correlations. If short-run movements in output were supply determined on the other hand, the inflation rate and the cyclical component of output should be negatively correlated.[10] Whether cyclical movements in output are determined primarily by fluctuations in demand or supply is an empirical question. We now proceed to examine the correlations between the rate of inflation and alternative measures of the cyclical component of output. Of course, for such correlations to be valid the rate of inflation must be a stationary variable. Appendix I examines the stationarity of the inflation rate in the postwar United States. The evidence suggests that the inflation rate is indeed stationary in our data.

Table 2 reports correlations between inflation and measures of the cyclical component of output under the assumption that long-run movements in output can either be described as functions of time or be picked up by the Hodrick-Prescott filter. The rate of inflation is measured as the first difference of the logarithm of the price level. The first panel reports the correlations between inflation and output that is detrended by removing a linear time trend. The cross-correlations are significantly positive at virtually all leads and lags. The second panel reports the correlations between the rate of inflation and output from which a segmented trend has been removed. The correlations in this panel are also virtually all positive. Finally, the third panel reports the correlations between the rate of inflation and output detrended using the Hodrick-Prescott filter. Again, all the correlations are positive. The results in Table 2 are in sharp contrast to the correlations reported in Table 1: in Table 2 virtually every reported correlation is positive, providing evidence that prices are in fact procyclical.

We now turn to an examination of the correlations between inflation and estimates of the cyclical component of output under the assumption that output is characterized by a unit root. The cyclical components of

[10] When output is supply determined, the interaction of the degree of price flexibility with the extent of persistence of the shocks affecting the economy will determine the pattern of the correlations. However, there will always be at least some negative correlations. Predicted correlations for reasonable parameter values are reported later in the paper.

Table 2. *Cross-Correlations of Inflation and Measures of Cyclical Output, Quarterly Data, 1947:2 to 1989:4*

| Lag | Cross-correlation | Standard error |
|---|---|---|
| Log difference of prices and detrended output | | |
| 4 | 0.33 | 0.07 |
| 3 | 0.29 | 0.07 |
| 2 | 0.25 | 0.07 |
| 1 | 0.24 | 0.07 |
| 0 | 0.21 | 0.07 |
| −1 | 0.19 | 0.08 |
| −2 | 0.16 | 0.08 |
| −3 | 0.09 | 0.08 |
| −4 | 0.04 | 0.08 |
| Log difference of prices and detrended output (trend break for output in 1973:1) | | |
| 4 | 0.16 | 0.08 |
| 3 | 0.14 | 0.08 |
| 2 | 0.13 | 0.08 |
| 1 | 0.13 | 0.08 |
| 0 | 0.13 | 0.08 |
| −1 | 0.12 | 0.08 |
| −2 | 0.09 | 0.08 |
| −3 | 0.03 | 0.08 |
| −4 | −0.04 | 0.08 |
| Log difference of prices and Hodrick-Prescott filtered output | | |
| 4 | 0.07 | 0.08 |
| 3 | 0.07 | 0.08 |
| 2 | 0.11 | 0.08 |
| 1 | 0.14 | 0.08 |
| 0 | 0.16 | 0.08 |
| −1 | 0.17 | 0.08 |
| −2 | 0.16 | 0.08 |
| −3 | 0.08 | 0.08 |
| −4 | 0.01 | 0.08 |

Note: Lag 4 indicates a correlation of inflation with the fourth lag of the transformed output series. A negative lag denotes a lead.

output are constructed using both the Beveridge-Nelson decomposition and the Blanchard-Quah methodology. Table 3 reports the correlations between the rate of inflation and the level of the stationary component of output obtained using the Beveridge-Nelson decomposition. The decomposition is implemented using the computational approach suggested by Cuddington and Winters (1987). It is well known that the

Table 3. *Cross-Correlations of Inflation and Cyclical Output Obtained Using the Beveridge-Nelson Decomposition, 1947:2 to 1989:4*

| Lag | Cross-correlation | Standard error |
|---|---|---|
| ARMA(1,1) | | |
| 4 | 0.07 | 0.08 |
| 3 | 0.06 | 0.08 |
| 2 | 0.02 | 0.08 |
| 1 | 0.00 | 0.08 |
| 0 | 0.01 | 0.08 |
| −1 | 0.03 | 0.08 |
| −2 | 0.07 | 0.08 |
| −3 | 0.18 | 0.08 |
| −4 | 0.20 | 0.08 |
| ARMA(2,2) | | |
| 4 | 0.04 | 0.08 |
| 3 | 0.00 | 0.08 |
| 2 | −0.03 | 0.08 |
| 1 | −0.03 | 0.08 |
| 0 | 0.02 | 0.08 |
| −1 | 0.04 | 0.08 |
| −2 | 0.07 | 0.08 |
| −3 | 0.15 | 0.08 |
| −4 | 0.15 | 0.08 |
| ARMA(5,1) | | |
| 4 | 0.22 | 0.08 |
| 3 | 0.23 | 0.08 |
| 2 | 0.25 | 0.08 |
| 1 | 0.22 | 0.08 |
| 0 | 0.20 | 0.08 |
| −1 | 0.19 | 0.08 |
| −2 | 0.17 | 0.08 |
| −3 | 0.12 | 0.08 |
| −4 | 0.06 | 0.08 |

Note: ARMA models were estimated for the log difference of output, allowing for a break in the growth rate of output in 1973:1 (by subtracting out mean growth rates of output for each of the two subperiods). The period over which correlations are computed depends on the order of the AR component. Lag 4 indicates the correlation of inflation with the fourth lag of transformed output. A negative lag indicates a lead.

decomposition is sensitive to the choice of ARIMA specification for the level of output (see, for example, Canova (1991)). Since there does not appear to be any generally accepted specification, various specifications were tried. The resulting correlations between inflation and the cyclical component were, however, found to be broadly similar to the three specifications reported in Table 3.

The first panel of Table 3 reports correlations between the rate of inflation and the cyclical component of output from a decomposition using an ARMA(1,1) specification for the first difference of output (that is, ARIMA(1,1,1) for the level of output). Most of the correlations are small and positive, except at the third and fourth leads, where the correlations are strongly positive. With an ARMA(2,2) specification, the results are similar, except for small negative correlations at the first and second lags. Using an ARMA(5,1) specification, we find strong positive correlations at virtually all lags and leads. In summary, the results vary between small positive correlations (with an occasional small negative correlation) and strong positive correlations.

Table 4 reports correlations between the rate of inflation and the level of the cyclical component of output obtained from alternative Blanchard-Quah decompositions of output. In the first panel of Table 4 the cyclical component of output was constructed to correspond to Blanchard and Quah's (1989) "base case": a bivariate system with the first difference of output and the level of unemployment was estimated allowing for a break in the growth rate of output in 1973:4 and a time trend in the unemployment rate.[11] The correlations are significantly positive at the third and fourth lags, gradually decline in magnitude as the lag length decreases, and take on small negative values at the third and fourth leads.[12] The second panel presents correlations between the rate of inflation and the cyclical component of output using the Blanchard-Quah methodology when the rate of inflation is used in place of the unemployment rate in the bivariate system. In principle, any variable that is stationary and is affected by the same shocks as output could be used in the Blanchard-Quah decomposition. Given the focus of this paper, it seems only natural to decompose output using the inflation rate.[13] The reported cross-

[11] The unemployment series is the civilian unemployment rate (seasonally adjusted) from the Data Resources Incorporated data bank.

[12] The decomposition is sensitive to assumptions about whether the rate of growth of output is modeled as stationary with or without a break in mean coincident with the 1973 oil shock and whether the unemployment rate is stationary in levels or around a time trend. Alternative treatments of the break and trend yielded mixed results but, even in cases where we found some negative correlations, they were small and insignificant. These results are available upon request.

[13] Shapiro and Watson (1988) use a decomposition technique similar to that of Blanchard and Quah and include inflation in their vector autoregression, although they include its first difference rather than its level since they argue that the inflation rate could be nonstationary. Our tests (Appendix I) lead us to conclude, however, that the inflation rate is stationary. In any case, when we implemented the Blanchard-Quah decomposition with the first differences of inflation, we still obtained positive correlations, although the magnitudes were smaller than those reported in the second panel of Table 4.

Table 4. *Cross-Correlations of Inflation and Cyclical Output Obtained Using the Blanchard-Quah (BQ) Decomposition, 1949:4 to 1989:4*

| Lag | Cross-correlation | Standard error |
|---|---|---|
| Cyclical output from bivariate BQ decomposition with output and unemployment | | |
| 4 | 0.17 | 0.08 |
| 3 | 0.16 | 0.08 |
| 2 | 0.13 | 0.08 |
| 1 | 0.11 | 0.08 |
| 0 | 0.09 | 0.08 |
| −1 | 0.07 | 0.08 |
| −2 | 0.05 | 0.08 |
| −3 | −0.01 | 0.08 |
| −4 | −0.05 | 0.08 |
| Cyclical output from bivariate BQ decomposition with output and inflation | | |
| 4 | 0.54 | 0.08 |
| 3 | 0.55 | 0.08 |
| 2 | 0.56 | 0.08 |
| 1 | 0.57 | 0.08 |
| 0 | 0.67 | 0.08 |
| −1 | 0.67 | 0.08 |
| −2 | 0.66 | 0.08 |
| −3 | 0.61 | 0.08 |
| −4 | 0.56 | 0.08 |

Note: In the top panel, the decomposition is similar to that presented as the base case in Blanchard and Quah (1989), with a break in the growth rate of output in 1973:4, and a time trend removed from the total civilian unemployment rate. The correlations in this table are computed from 1949:4 since unemployment data is available only from 1948:1 and six lags are used in estimating the vector autoregression. Lag 4 indicates the correlation of inflation with the fourth lag of cyclical output. A negative lag indicates a lead.

correlations between inflation and the cyclical component of output are now all very strongly positive.

The results in Tables 3 and 4 suggest that, under the assumption that output is characterized by a unit root, there is typically a positive correlation between the rate of inflation and cyclical movements in output. The magnitude of the correlations and, in a small number of cases, the sign, is, however, sensitive to the methods used to decompose output into permanent and stationary components. Significantly positive correlations were found for a Beveridge-Nelson decomposition from an ARMA(5,1) process fitted to the first difference of output and from a Blanchard-Quah decomposition of output from a bivariate system with the first difference of output and the rate of inflation.

## II. Output and Price Correlations in Two Simple Macroeconomic Models

This section illustrates the effects of applying alternative transformations to prices and output by examining two simple stylized macroeconomic models: a sticky-price demand-driven model and a flexible-price supply-driven model. The models are solved for the impulse response functions of prices and output to innovations in demand and supply. The implications of applying alternative transformations to the series are discussed. Stochastic simulations of the models are then conducted, and the cross-correlations between alternative transformations of the generated price and output data are examined.

Consider a stylized economy where the level of output is determined by the demand for it. The demand for output is assumed to be a positive function of the level of real money balances and a demand shift term, so that

$$y_t = y_t^d = M_t - P_t + D_t, \tag{1}$$

where $y_t$ denotes the logarithm of output at time $t$; the superscript $d$ denotes demand; $M_t$ is the logarithm of the nominal money stock; $P_t$ represents the logarithm of the aggregate price level; and $D_t$ is the demand-shift term. Equation (1) can be motivated by a simple quantity theory equation.[14] The money supply is assumed to grow at a perfectly predictable constant rate, $\mu$, so that

$$M_t = M_{t-1} + \mu, \tag{2}$$

and the demand-shift term, $D_t$, is specified as an AR(1) process

$$D_t = \rho D_{t-1} + \epsilon_t, \qquad 0 \leq \rho \leq 1 \tag{3}$$

where $E_t(\epsilon_{t+v}) = 0$, for all $v = 1, 2, \ldots$, and $\mathrm{var}(\epsilon) = \sigma_\epsilon^2$. $E_t$ denotes the mathematical expectations operator conditional on information available at time $t$. Note from equation (1) that $D_t$ can be interpreted as representing nominal or real shocks. Moreover, these shocks could be permanent ($\rho = 1$) or temporary ($\rho \leq 1$) since even permanent shocks to demand will have a temporary effect on output as prices adjust. Both cases are discussed below.

The logarithm of the long-run level of output, which can be thought of as a measure of capacity or natural level of output, and which we refer

[14] Alternatively, it can be viewed as a (linear function of the) reduced-form solution to a (log-linear) simultaneous IS-LM system, for any price level given by history.

to as the supply of output, is assumed for simplicity to be a deterministic function of time:

$$y_t^s = a + bt. \tag{4}$$

By construction, there are no supply shocks in this stylized economy. It is worth emphasizing at the outset that all of the points we make in the context of this example are equally applicable to the case where output has a unit root—that is, a stochastic trend. Since we are interested in the cyclical, or temporary, component of output, it matters little what process describes the long-run, or permanent, movements. Assuming a time trend, however, has the practical advantage of permitting the use of a simple detrending procedure to obtain the cyclical component of output. An example of a case where output has a unit root is presented in Appendix II.

The price level is assumed to be sticky in that it is completely predetermined at a point in time and adjusts only slowly to equilibrate the goods market. The dynamic adjustment process for the price level is assumed to be given by a form of the Barro-Grossman rule, which posits adjustment to be a function of both the change in equilibrium prices and the extent of disequilibrium in the goods market. The particular rule employed is a version of that proposed by Mussa (1981a, 1981b):[15]

$$P_{t+1} - P_t = \overline{\Pi}_{t+1} + \beta[y_t^d - y_t^s], \qquad \text{where } 0 < \beta < 1. \tag{5}$$

In equation (5), $\overline{\Pi}_{t+1}$ denotes the time $t$ expectation of the long-run equilibrium inflation rate, which is formally defined as

$$\overline{\Pi}_{t+1} = \lim_{v \to \infty} E_t[\overline{P}_{t+1+v} - \overline{P}_{t+v}], \tag{6}$$

where $\overline{P}_t$ denotes the equilibrium or flexible-price level in period $t$.

The equilibrium or flexible-price level, $\overline{P}_t$, is defined as the price level that equilibrates the goods market in each period, that is $y_t^d = y_t^s$, and therefore

$$\overline{P}_t = M_t + D_t - y_t^s = \overline{P}_{t-1} + (\mu - b) + D_t - D_{t-1}. \tag{7}$$

Substituting equations (6) and (7) into (5), the solution for the price level can be written as

$$P_t = (1 - \beta)P_{t-1} + \beta\overline{P}_{t-1} + (\mu - b). \tag{8}$$

The aggregate price level is, therefore, a weighted average of the previous period's price level and the previous period's flexible-price level, adjusted

[15] Obstfeld and Rogoff (1984) discuss the appropriateness of alternative sticky-goods-price adjustment rules.

for the drift in prices given by $(\mu - b)$. The first term represents the inertia in the price level while the second term represents the stochastic long-run equilibrium toward which the price level adjusts.

We now define $\hat{P}_t$ to be the extent of disequilibrium in the goods market measured as the difference between the flexible-price level and the actual price level. Then, substituting in the solutions for the flexible-price level and the actual price level from equations (7) and (8),

$$\hat{P}_t = \overline{P}_t - P_t = (1 - \beta)\hat{P}_{t-1} + \epsilon_t - (1 - \rho)D_{t-1}, \tag{9}$$

so that disequilibrium in the goods market equals (i) a fraction, $(1 - \beta)$, of the disequilibrium from the previous period that is not dissipated by price changes; (ii) the innovation in the demand-shift term, $\epsilon_t$, which adds one-for-one to the existing disequilibrium since prices only begin to adjust with a lag; (iii) less a fraction, $(1 - \rho)$, of the shock since the shock itself dissipates over time and in that sense restores equilibrium. The extent of disequilibrium is, therefore, a stationary ARMA process that reverts to its mean value of zero over time as prices adjust to a past shock or as the shock that created the disequilibrium dissipates.

Output is determined by demand. Adding and subtracting the supply of output from both sides of equation (1), and employing the definition of disequilibrium in the goods market measured as $\hat{P}_t$ in equation (9), the solution for output can be written as

$$y_t = a + bt + \hat{P}_t. \tag{10}$$

Output is, therefore, stationary around a time trend defined by the supply of output. Consequently, the appropriate method of recovering the cyclical component of output is the standard detrending procedure of regressing the variable on a constant and a time trend. Though the particular dynamics of output around its trend will depend on the persistence and permanence of innovations in the demand-shift term, output will always be stationary around a deterministic trend.

It is useful to write the price level from equation (9) as

$$P_t = \overline{P}_t - \hat{P}_t, \tag{11}$$

so that the price level can then be viewed as the sum (difference) of a nonstationary component, the flexible-price level, and a stationary component representing the extent of disequilibrium. The price level will, therefore, exhibit stationary deviations described by an ARMA process, representing the evolution of disequilibrium, $\hat{P}_t$, around a nonstationary or permanent component representing the evolution of the flexible-price level. The particular form of nonstationarity of the flexible-price level will depend on whether the shocks affecting demand are temporary or permanent. Using equations (11) and (7), the price level can be expressed as a function of time:

$$P_t = (M_0 - a) + (\mu - b)t + \sum_{i=0}^{\infty} \rho^i \epsilon_{t-i} - \hat{P}_t. \tag{12}$$

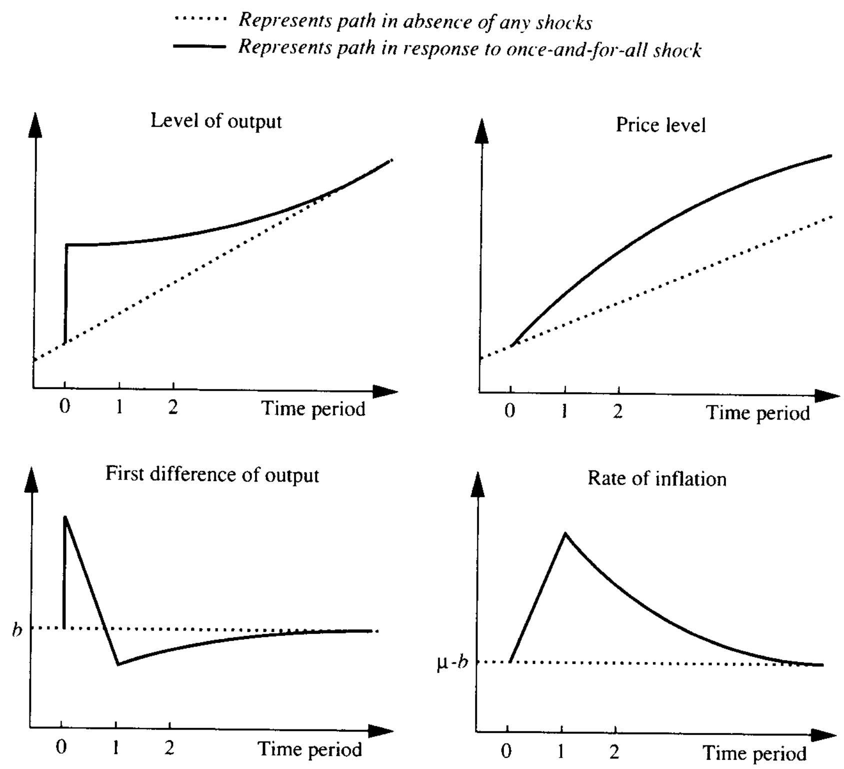
Note that, unlike the case of output that was described by movements of $\hat{P}_t$ around a time trend in equation (10), there are now two terms describing movements of the price level around a time trend. The first represents innovations that affect the flexible-price level; the second term, $\hat{P}_t$, represents the extent of disequilibrium. While $\hat{P}_t$ is always stationary, whether or not the two components are jointly stationary will depend on whether $\rho$ is less than 1 or whether $\rho$ is equal to 1.

Consider first the case where all shocks to demand are permanent and $\rho = 1$. Then the summation term in equation (12) implies that the price level is nonstationary around a time trend as the entire path of the price level will shift in response to an innovation, $\epsilon_t$, in the demand-shift term. The price level will, therefore, be characterized by unit-root nonstationarity. When $\rho = 1$ the extent of disequilibrium, $\hat{P}_t$, will, from equation (9), be a pure AR(1) process. Figure 1 plots the implied impulse responses of output and the price level to a once-and-for-all positive innovation in the demand-shift term. Output rises on impact in period 0 and gradually returns to trend. The price level begins to rise (with a lag) faster than implied by its drift or trend and then settles at a permanently higher level.

As noted earlier, by construction the appropriate method of recovering the cyclical component of output in our example is to regress output on a constant and a time trend. Regressing the price level on a constant and a time trend, however, and then examining the cross-correlations of detrended output and prices will yield unpredictable results. Estimating equation (12) when $\rho = 1$ on the assumption that the price level returns to a fixed path will then artificially attribute some of the observations to be below and some to be above an estimated trend line.[16] Using deviations of these observations from the trend line as a cyclical measure of price movements would then yield spurious results. Figure 1 also plots the impulse responses of the first difference of output and the rate of inflation. First differencing output yields a constant (equal to the long-run growth rate) and the change in the cyclical component of output. While output jumps up on impact implying a positive change in output, it then

[16] Note that this trend line would not correspond to the dashed line in the panel for the price level in Figure 1.

Figure 1. *Impulse Response Functions in Demand-Determined Model When Shocks Are Permanent*




starts to return to its trend level implying a negative change in output.[17] This reversal will induce some negative correlations between the first difference of output and the rate of inflation. The important point to note is that, except for the period in which the shock occurs and the change in output is positive, the first difference of output is an erroneous indicator of the level of the cyclical component of output. Use of the first difference of output as an indicator of cyclical movements in output would imply that, except for the first period, output is below its average value when in fact it is above.

[17] Note that in this and the subsequent discussion, movements in all variables are referred to as relative to their trend or, in the case of stationary variables, relative to their average values.

The rate of inflation can, by first differencing equation (8), be written as

$$\Pi_t = P_t - P_{t-1} = (\mu - b) + \beta \hat{P}_{t-1}. \tag{13}$$

The impulse response of inflation around its average value will, therefore, be deterministically related to the level of the cyclical component of output lagged one period, $\hat{P}_{t-1}$. Since the cyclical component of output is an AR(1) process and hence autocorrelated, the rate of inflation will be correlated with it at several lags and leads. Figure 1 shows in fact that this is the only robust positive cross-correlation that should be expected from the three measures discussed. It is worth noting that equation (13) does not depend on the nature of the shocks.

Consider now the case of temporary shocks to demand—that is, when $\rho < 1$ in equation (3). To take the simplest case first, let $\rho = 0$. In this case shocks are completely temporary. As Figure 2 shows, upon impact of a shock in period 0, output will, as before, jump above its long-run trend level by the amount of the shock, creating excess demand. In period 1, the shock disappears. However, since prices adjust with a lag to excess demand, the price level will rise in period 1. This translates into a net negative effect on aggregate demand in period 1 and will result in a decline in output relative to trend. Subsequently, output will rise and return gradually to its long-run trend level while the price level will fall. The plots in Figure 2 show that the impulse responses are not monotonic and can in fact be fairly complicated. The cross-correlations between both detrended prices and output and the first differences of the series will, therefore, be mixed in sign. Note that the response functions that are similar are those of inflation lagged one period and detrended output.

Consider now the effect of a nonzero $\rho$. The stochastic simulations below cover a range of values of $\rho$ so here we briefly discuss an intuitive generalization of the $\rho = 0$ case. Essentially, as $\rho$ increases from zero, the shock, although it remains temporary, becomes more persistent. The patterns of the impulse response functions presented in Figure 2 will remain the same. Instead of the sharp changes in sign, however, there will be smoother and more gradual changes in sign.

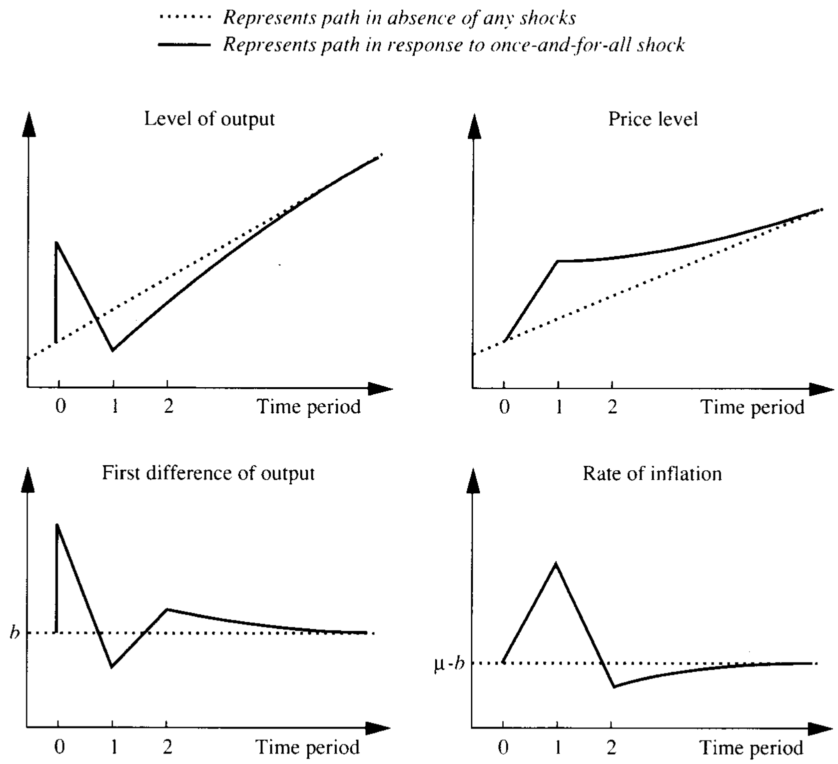
We turn now to an economy where the level of output is supply determined. It is assumed again that long-run movements in output are described by a time trend, so that

$$y_t = y_t^s = a + bt + S_t, \tag{14}$$

where $S_t$ is a supply-shift term that represents the cyclical component of output and is, by definition, stationary. The process generating $S_t$ is assumed to be an AR(1) process:

$$S_t = \gamma S_{t-1} + \Omega_t, \qquad 0 \leq \gamma < 1, \tag{15}$$

Figure 2. *Impulse Response Functions in Demand-Determined Model When Shocks Are Temporary*




and where $E_t(\Omega_{t+v}) = 0$, for all $v = 1, 2, \ldots,$ and $\operatorname{var}(\Omega) = \sigma_\Omega^2$. As before, aggregate demand is assumed to be a positive function of the level of real money balances

$$y_t^d = M_t - P_t, \tag{16}$$

and the nominal money supply is assumed to grow at a constant rate, $\mu$. It is assumed that prices are flexible.[18] The price level is defined by the equality of aggregate demand and supply:

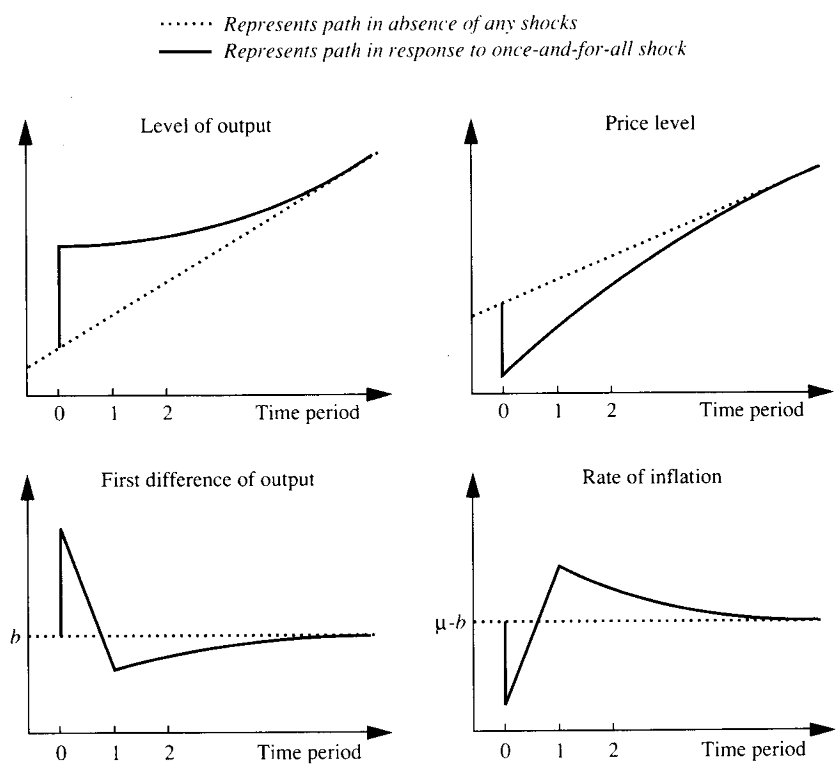
$$y_t^d = y_t^s, \quad \text{or} \quad \overline{P}_t = M_t - y_t^s = (M_0 - a) + (\mu - b)t - S_t. \tag{17}$$

[18] If output is supply determined (and independent of the price level as it is here), price stickiness will not affect the dynamic behavior of output. It will affect only the dynamics of the price level.

In this case, both output and prices are stationary around a time trend. Moreover, the cyclical, or detrended, component of prices is simply the negative of the detrended component of output (see equation (14)).

The impulse responses of output and prices are plotted in Figure 3. Detrended output and prices will be perfectly negatively correlated contemporaneously and, since the supply-shift term is positively autocorrelated, the cross-correlations of the two series over time will also be negative. Again, the first difference of output will be an erroneous indicator of cyclical movements in output. The rate of inflation falls below its long-run value on impact of the shock, as the price level falls; it then rises above its long-run value as the price level begins to rise and then gradually falls to its long-run value. A key implication of the impulse response of inflation is that there will be a negative contemporaneous

Figure 3. *Impulse Response Functions in Supply-Determined Model When Shocks Are Temporary*



correlation between the innovation in the cyclical component of output and the rate of inflation.

## III. Stochastic Simulations

The two models defined by equations (1)–(6) and (14)–(17) were simulated using numerical values for the parameters $a$, $b$, $\mu$, $\beta$, $\rho$, and $\gamma$. The simulations were carried out using the random number generator in the econometric software package RATS. The parameters were chosen to be consistent with quarterly data so that a time period in the simulations should be interpreted as a quarter. The standard deviations of the innovations in the (logarithms of the) demand-shift and supply-shift terms were both set at 1 percent.

The constant $a$ in the supply of output equation was set at zero, while the long-run rate of growth of output, determined by the parameter $b$, was set to imply an annual growth rate of 3 percent. The parameter $\mu$ was set so that the money supply was expected to grow at an annual rate of 7 percent. These assumptions imply an average or long-run annual inflation rate of 4 percent.

For the demand-driven model, a key parameter is the degree of price flexibility as measured by the parameter $\beta$. Rather than perform the simulations on a wide range of alternative parameter values, a representative value of 0.05 on a quarterly basis for the United States was chosen. This value is consistent with estimates of price stickiness made for the United States by Taylor (1980)[19] and Rotemberg (1982);[20] it implies that about 20 percent of the gap between actual prices and their flexible equilibrium solution is made up in one year. To allow for both permanent and temporary demand shocks, simulations were performed with $\rho = 1$ (all shocks permanent) and values of $\rho$ less than 1 (all shocks are temporary). In the case of the supply-driven model, $\gamma$ was set at 0.9. This value compares with the first-order serial correlation of 0.95 assumed

[19] Taylor (1980) estimates the response of "new" wage contracts to excess demand as 0.087 on a quarterly basis. Since in his framework only a subset of wage-price setters revises its prices in each period, our value of 0.05 for the responsiveness of the aggregate price level seems appropriate.

[20] Rotemberg (1982) jointly estimates a system of price, output, and money equations, with a reduced-form equation for the general price level very similar to the solution obtained in equation (8) once the flexible-price level is substituted out for and expressed in terms of the money supply. His preferred estimates for the coefficient on the lagged price level vary from 0.92 to 0.95; this coefficient corresponds to $(1 - \beta)$ in our system, suggesting a value for $\beta$ between 0.08 and 0.05. For further details, see Rotemberg (1982) and Chadha (1989).

by Kydland and Prescott (1982) in the total factor productivity shocks affecting the supply of output.

A relatively large sample size of 400 observations (100 years) was chosen, and three versions of the demand-driven model and two versions of the supply-driven model were each simulated 100 times. A simulation is referred to as a "run." Three sets of cross-correlations of prices and output were constructed for the data generated in each run of the model: between detrended prices and output; between the rate of inflation and the first difference of output; and between the rate of inflation and detrended output. The results for each measure of the co-movement between prices and output across the runs are summarized in Tables 5 and 6. To preserve any systematic pattern in the correlations implied by the dynamics of the model, the vectors of correlations reported correspond to particular runs. They were ordered on the basis of the magnitude of the cross-correlation at lag 1 for the demand-driven model, since the model predicts that the correlation between price changes and the cyclical component of output will be largest at the first lag. In the case of the supply-driven model, when prices are flexible the vectors were ordered by the contemporaneous correlation. Tables 5 and 6 also report the standard deviations of the elements of the correlation vectors across the 100 runs of each version of the models.

For the demand-driven model, when all shocks are permanent and $\rho = 1$ (first set of columns of Table 5), while output is stationary around a time trend, the price level contains a unit root. The cross-correlations of the detrended series from the generated data will in general, therefore, be spurious. This is reflected in the reported vectors of correlations, which range from having consistently large negative entries to consistently large positive entries. All vectors of cross-correlations between the first differences of the series display the same pattern.[21] While correlations at all lags are positive, the contemporaneous and all led correlations are negative. The observed pattern is, as noted above, a consequence of first differencing the output series; this yields the first difference of the cyclical component of output rather than its level and induces the negative contemporaneous and led correlations. Finally, the cross-correlations of the rate of inflation with detrended output are examined: all elements of all correlation vectors are positive.

In the demand-driven model, when shocks are temporary and $\rho = 0.9$ or $\rho = 0.5$ (second and third sets of columns in Table 5), comparisons

[21] They are also rather small in absolute value.

between detrended data and between first-differenced data typically yield a similar pattern. As one would expect from the discussion of temporary demand shocks, there are both negative and positive correlations: lagged correlations are typically positive while led and contemporaneous correlations are typically negative. The only procedure that produces relatively robust positive cross-correlations is a comparison between the rate of inflation and detrended output.

Clearly then, examining the cross-correlations of either detrended or first-differenced data to determine the co-movement of prices with the cyclical component of output can, when output is demand determined, lead to spurious conclusions. The simulations have shown that in an economy where output is demand determined and prices are procyclical by construction, the two measures can easily yield negative correlations. A more accurate measure of the co-movement is obtained by examining the correlations between the rate of inflation and the level of the cyclical component of output.

In the supply-driven model, where the cyclical component of output results entirely from shocks to supply, both output and the price level are stationary around a time trend. When prices are flexible (first set of columns of Table 6), all reported vectors of cross-correlations between the detrended values of the two series are consistently and significantly negative, accurately reflecting the countercyclical behavior of prices. The first differences of the series are typically found to be negligibly positively correlated except for the contemporaneous correlation, which is almost always negative unity. The cross-correlations between inflation and detrended output yield both positive and negative values: all lagged correlations are positive while the contemporaneous and all led correlations are negative. Clearly then, when output is supply determined and prices are flexible, the measure that most accurately reflects the countercyclical behavior of prices is the negative correlation between the levels of cyclical components of output and prices—here represented by the detrended values of the series. This negative correlation, however, carries over, for the contemporaneous and led values, to a comparison between the rate of inflation and detrended output. The switch in sign at the contemporaneous value occurs because of the assumption of complete price flexibility.

Table 6 (second set of columns) also reports the results of assuming that prices are sticky as in the demand-driven model with $\beta$, the price responsiveness parameter, set at the same value 0.05. There is, then, a robust negative correlation between both the detrended values of the two series and between inflation and detrended output.

Table 5. *Summary of Cross-Correlations Obtained from Stochastic Simulations of Demand-Determined Output Model*

| | $\rho = 1$ | | | | $\rho = 0.9$ | | | | $\rho = 0.5$ | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Lag | Minimum | Median | Maximum | Standard deviation | Minimum | Median | Maximum | Standard deviation | Minimum | Median | Maximum | Standard deviation |
| Detrended price level and detrended output | | | | | | | | | | | | |
| 4 | −0.34 | 0.11 | 0.37 | 0.12 | 0.17 | 0.32 | 0.34 | 0.05 | 0.20 | 0.28 | 0.29 | 0.04 |
| 3 | −0.36 | 0.08 | 0.35 | 0.12 | 0.13 | 0.24 | 0.31 | 0.04 | 0.18 | 0.26 | 0.31 | 0.04 |
| 2 | −0.38 | 0.05 | 0.34 | 0.12 | 0.08 | 0.16 | 0.26 | 0.04 | 0.14 | 0.21 | 0.28 | 0.03 |
| 1 | −0.40 | 0.02 | 0.32 | 0.12 | 0.01 | 0.05 | 0.19 | 0.04 | 0.07 | 0.11 | 0.19 | 0.02 |
| 0 | −0.43 | −0.01 | 0.30 | 0.12 | −0.08 | −0.06 | 0.11 | 0.04 | −0.08 | −0.11 | −0.04 | 0.02 |
| −1 | −0.45 | −0.05 | 0.28 | 0.12 | −0.15 | −0.16 | 0.03 | 0.04 | −0.15 | −0.21 | −0.13 | 0.03 |
| −2 | −0.48 | −0.08 | 0.26 | 0.12 | −0.21 | −0.25 | −0.04 | 0.05 | −0.19 | −0.26 | −0.17 | 0.04 |
| −3 | −0.50 | −0.12 | 0.24 | 0.12 | −0.24 | −0.33 | −0.10 | 0.06 | −0.21 | −0.28 | −0.17 | 0.04 |
| −4 | −0.52 | −0.15 | 0.22 | 0.13 | −0.27 | −0.39 | −0.15 | 0.06 | −0.21 | −0.27 | −0.16 | 0.04 |
| First difference of price level and first difference of output | | | | | | | | | | | | |
| 4 | 0.13 | 0.14 | 0.17 | 0.03 | 0.27 | 0.16 | 0.16 | 0.03 | 0.13 | 0.06 | 0.06 | 0.04 |
| 3 | 0.12 | 0.16 | 0.20 | 0.03 | 0.26 | 0.17 | 0.21 | 0.03 | 0.17 | 0.09 | 0.01 | 0.04 |
| 2 | 0.10 | 0.17 | 0.23 | 0.03 | 0.24 | 0.22 | 0.31 | 0.03 | 0.25 | 0.29 | 0.22 | 0.04 |
| 1 | 0.13 | 0.17 | 0.24 | 0.02 | 0.22 | 0.28 | 0.34 | 0.03 | 0.44 | 0.52 | 0.57 | 0.02 |
| 0 | −0.13 | −0.15 | −0.23 | 0.03 | −0.21 | −0.29 | −0.34 | 0.03 | −0.45 | −0.52 | −0.57 | 0.02 |
| −1 | −0.10 | −0.15 | −0.22 | 0.03 | −0.23 | −0.23 | −0.31 | 0.03 | −0.25 | −0.29 | −0.21 | 0.04 |
| −2 | −0.12 | −0.14 | −0.19 | 0.03 | −0.25 | −0.18 | −0.21 | 0.03 | −0.17 | −0.10 | −0.01 | 0.04 |
| −3 | −0.12 | −0.12 | −0.16 | 0.03 | −0.26 | −0.16 | −0.16 | 0.03 | −0.13 | −0.07 | −0.06 | 0.04 |
| −4 | −0.12 | −0.12 | −0.17 | 0.03 | −0.24 | −0.17 | −0.14 | 0.03 | −0.07 | 0.02 | −0.03 | 0.04 |

| | | | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| First difference of price level and detrended output | | | | | | | | | | | | |
| 4 | 0.45 | 0.83 | 0.85 | 0.08 | 0.60 | 0.51 | 0.62 | 0.07 | 0.04 | 0.11 | 0.07 | 0.06 |
| 3 | 0.52 | 0.88 | 0.89 | 0.07 | 0.71 | 0.65 | 0.71 | 0.05 | 0.14 | 0.23 | 0.15 | 0.06 |
| 2 | 0.59 | 0.93 | 0.95 | 0.06 | 0.84 | 0.80 | 0.83 | 0.03 | 0.44 | 0.49 | 0.43 | 0.04 |
| 1 | 0.67 | 0.98 | 1.00 | 0.06 | 0.98 | 0.99 | 1.00 | 0.00 | 1.00 | 1.00 | 1.00 | 0.02 |
| 0 | 0.59 | 0.93 | 0.95 | 0.06 | 0.84 | 0.80 | 0.83 | 0.03 | 0.44 | 0.49 | 0.43 | 0.04 |
| −1 | 0.53 | 0.88 | 0.90 | 0.07 | 0.72 | 0.65 | 0.71 | 0.05 | 0.14 | 0.23 | 0.15 | 0.06 |
| −2 | 0.46 | 0.84 | 0.85 | 0.08 | 0.60 | 0.52 | 0.62 | 0.07 | 0.04 | 0.11 | 0.07 | 0.06 |
| −3 | 0.42 | 0.79 | 0.81 | 0.08 | 0.49 | 0.41 | 0.51 | 0.08 | −0.02 | −0.03 | −0.05 | 0.06 |
| −4 | 0.38 | 0.74 | 0.78 | 0.09 | 0.41 | 0.30 | 0.43 | 0.09 | 0.01 | 0.00 | −0.05 | 0.05 |

Note: One hundred runs of each version of the model were simulated. The sample size was set at 400 quarters. Each reported cross-correlation vector corresponds to a particular run. Runs were ordered on the basis of the magnitude of the correlation at lag 1 to obtain the "minimum," "maximum," and "median" vectors. The standard deviation is that of elements of the correlation vectors across the 100 runs of each model.

Table 6. *Summary of Cross-Correlations Obtained from Stochastic Simulations of Supply-Determined Output Model*

| | Flexible prices | | | | Sticky prices | | | |
|---|---|---|---|---|---|---|---|---|
| Lag | Minimum | Median | Maximum | Standard deviation | Minimum | Median | Maximum | Standard deviation |
| Detrended price level and detrended output | | | | | | | | |
| 4 | −0.70 | −0.58 | −0.60 | 0.08 | −0.80 | −0.68 | −0.64 | 0.04 |
| 3 | −0.75 | −0.66 | −0.71 | 0.07 | −0.79 | −0.65 | −0.59 | 0.04 |
| 2 | −0.82 | −0.76 | −0.80 | 0.05 | −0.77 | −0.61 | −0.51 | 0.05 |
| 1 | −0.89 | −0.88 | −0.89 | 0.03 | −0.75 | −0.55 | −0.42 | 0.06 |
| 0 | −1.00 | −1.00 | −1.00 | 0.00 | −0.71 | −0.48 | −0.29 | 0.08 |
| −1 | −0.89 | −0.88 | −0.89 | 0.03 | −0.68 | −0.41 | −0.18 | 0.09 |
| −2 | −0.82 | −0.76 | −0.80 | 0.05 | −0.66 | −0.36 | −0.09 | 0.11 |
| −3 | −0.75 | −0.66 | −0.71 | 0.07 | −0.64 | −0.31 | 0.00 | 0.12 |
| −4 | −0.70 | −0.58 | −0.60 | 0.08 | −0.62 | −0.27 | 0.07 | 0.13 |
| First difference of price level and first difference of output | | | | | | | | |
| 4 | 0.02 | 0.00 | 0.02 | 0.05 | −0.19 | −0.26 | −0.29 | 0.03 |
| 3 | 0.07 | 0.12 | −0.02 | 0.05 | −0.24 | −0.28 | −0.29 | 0.03 |
| 2 | 0.14 | 0.02 | 0.03 | 0.05 | −0.36 | −0.34 | −0.34 | 0.03 |
| 1 | 0.04 | 0.17 | 0.11 | 0.05 | −0.41 | −0.36 | −0.33 | 0.02 |
| 0 | −0.99 | −0.99 | −1.00 | 0.00 | 0.27 | 0.20 | 0.11 | 0.03 |
| −1 | 0.04 | 0.17 | 0.11 | 0.05 | 0.26 | 0.20 | 0.13 | 0.04 |
| −2 | 0.14 | 0.01 | 0.04 | 0.05 | 0.17 | 0.17 | 0.12 | 0.04 |
| −3 | 0.07 | 0.12 | −0.01 | 0.05 | 0.13 | 0.17 | 0.13 | 0.03 |
| −4 | 0.02 | 0.00 | 0.02 | 0.05 | 0.12 | 0.12 | 0.15 | 0.03 |
| First difference of price level and detrended output | | | | | | | | |
| 4 | 0.15 | 0.16 | 0.09 | 0.03 | −0.41 | −0.44 | −0.28 | 0.06 |
| 3 | 0.19 | 0.20 | 0.12 | 0.03 | −0.56 | −0.56 | −0.40 | 0.05 |
| 2 | 0.28 | 0.24 | 0.15 | 0.03 | −0.73 | −0.70 | −0.54 | 0.04 |
| 1 | 0.31 | 0.23 | 0.18 | 0.03 | −0.93 | −0.87 | −0.71 | 0.04 |
| 0 | −0.30 | −0.25 | −0.20 | 0.03 | −0.82 | −0.79 | −0.60 | 0.04 |
| −1 | −0.28 | −0.24 | −0.17 | 0.03 | −0.73 | −0.72 | −0.51 | 0.05 |
| −2 | −0.19 | −0.20 | −0.13 | 0.03 | −0.64 | −0.67 | −0.44 | 0.06 |
| −3 | −0.15 | −0.16 | −0.10 | 0.03 | −0.56 | −0.61 | −0.38 | 0.08 |
| −4 | −0.13 | −0.16 | −0.10 | 0.03 | −0.50 | −0.55 | −0.34 | 0.09 |

Note: One hundred runs of each version of the model were simulated. The sample size was set at 400 quarters. Each reported cross-correlation vector corresponds to a particular run. The 100 runs for each version of the model were ordered on the basis of the magnitude of the contemporaneous correlation when prices are flexible and on the magnitude of the correlation at lag 1 when prices are sticky, to obtain the "minimum," "maximum," and "median" vectors. The standard deviation is that of elements of the correlation vectors across the 100 runs of each model.

## IV. Conclusion

This paper has examined the co-movement of prices with the cyclical component of output. We have argued that the popular and seemingly innocuous procedure of applying the same transformation to the levels of both output and prices to render them stationary, and then examining the cross-correlations of the resulting series, can be misleading. We have argued that a more appropriate procedure is to examine the correlations between the rate of inflation and the level of the cyclical component of output. If cyclical movements in output result primarily from movements in demand, then the cross-correlations should be positive. If they result primarily from movements in supply, the cross-correlations should be negative.

In examining postwar U.S. data, we first showed that the cross-correlations between similarly transformed price and output data are consistently and often strongly negative, as reported recently by a number of authors (Kydland and Prescott (1990), Backus and Kehoe (1992), Cooley and Ohanian (1991), and Smith (1992)). This finding has been interpreted by these authors as evidence of countercyclical price behavior. We then showed that the rate of inflation is consistently and often strongly positively correlated with various measures of the cyclical component of output estimated under alternative assumptions of the long-run behavior of output. These latter results are consistent with prices having been procyclical in the postwar United States and the view that temporary movements in output are primarily associated with movements in demand.

Two simple macroeconomic models have been presented which help reconcile some of the conflicting findings. In particular, stochastic simulations of the two models were used to illustrate that similarly transformed price and output data could easily yield negative correlations even when prices were, by construction, procyclical. For both models, correlations between the rate of inflation and the cyclical component of output were shown to accurately reflect the cyclical behavior of prices.

## APPENDIX I

### Stationarity of the Rate of Inflation

This appendix examines the stationarity of the inflation rate in postwar U.S. data at a quarterly frequency. For purposes of comparison, all test statistics estimated for the rate of inflation are also reported for the first difference of output. Table A1 reports the results of standard Dickey-Fuller and Augmented

Table A1. *Dickey-Fuller (DF) and Augmented Dickey-Fuller (ADF) Tests for a Unit Root, 1947:2 to 1989:4*

| | DF statistic | $Q$-statistic significance level | ADF statistic (one lag) | $Q$-statistic | ADF statistic (four lags) | $Q$-statistic significance level |
|---|---|---|---|---|---|---|
| Inflation rate | −6.02 | 0.06 | −3.99 | 0.97 | −3.16 | 0.88 |
| First difference of output | −8.79 | 0.74 | −6.34 | 0.72 | −5.87 | 0.95 |

Note: The regressions include a constant. The critical value at the 5 percent significance level is −2.89 from Fuller (1976). The $Q$-statistic provides a test of 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.

Dickey-Fuller (ADF) regressions for these variables. The $t$-statistics from the Dickey-Fuller regressions and from Augmented Dickey-Fuller regressions with one and four lags are all significant at the 5 percent level, suggesting that the inflation rate and the first difference of output are stationary. 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 uncorrelated. For each regression reported in the table, the $Q$-statistic has a significance level greater than 5 percent, indicating that the hypothesis that the residuals from these regressions are white noise cannot be rejected at the 5 percent level of significance. When the Augmented Dickey-Fuller regressions allow for four lags, the significance level of the $Q$-statistic rises to almost 90 percent for the inflation rate and 95 percent for the first difference of output suggesting that allowing for a maximum lag length of four in the ADF regressions is sufficient for both variables.

As a complement to the above tests, we also employ a stationarity test that is based on estimating the ratio of $2\Pi$ times the spectral density of the first difference of a series at frequency zero, to the variance of its first difference (see Huizinga (1987)). 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 Dickey-Fuller and Augmented Dickey-Fuller 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 A2 reports estimates of the normalized density for the inflation rate and for the first difference of output. For both these series, the estimated normalized density declines as the number of lagged autocorrelations employed increases; the normalized density approaches zero at the maximum number of lags employed (168): 0.02 for the rate of inflation and 0.01 for the first difference of output. These results support the hypothesis that the inflation rate and the first

Table A2. *Estimates of Normalized Spectral Density Function, 1947:2 to 1989:4*

| Number of lags | Inflation rate | First difference of output |
|---|---|---|
| 6 | 0.26 | 0.23 |
| 12 | 0.15 | 0.14 |
| 24 | 0.10 | 0.07 |
| 48 | 0.07 | 0.03 |
| 96 | 0.04 | 0.02 |
| 168 | 0.02 | 0.01 |

Note: The first column refers to the number of lagged autocorrelations used in constructing the estimate of the normalized spectral density function.

difference of output are stationary series. Although the fact that the normalized density approaches zero does not establish that the series are stationary, the results suggest that, even if there exist permanent stochastic components in the rate of inflation and the first difference of output, these components are very small.

## APPENDIX II

### Macroeconomic Model with a Stochastic Trend in Output

This appendix extends the demand-determined model of output where long-run movements in output are described by a deterministic time trend, to allow for a stochastic trend in output. The objective is to show that all of the arguments made earlier regarding the appropriate methods of determining the cyclical behavior of prices are equally applicable to the case where output has a stochastic trend.

Consider an economy where output is, as before, demand determined, and the demand for output is described by equations (1)–(3) in the text. The supply of output is assumed to be described by a random walk, so that equation (4) is replaced by[22]

$$y_t^s = y_{t-1}^s + b + \omega_t,$$
$$\text{where } E_t(\omega_{t+v}) = 0, \text{ for } v = 1, 2, \ldots, \text{ and } \operatorname{var}(\omega) = \sigma_\omega^2. \qquad \text{(A1)}$$

For simplicity, we limit the analysis to the case where all shocks to demand are permanent, that is, $\rho = 1.0$ in equation (3), and the shocks to demand and supply are assumed to be uncorrelated. Then, equating demand and supply, the flexible-price level can be solved for as

$$\overline{P}_{t+1} = \overline{P}_t + (\mu - b) + \epsilon_{t+1} - \omega_{t+1}. \qquad \text{(A2)}$$

[22] This particular form of stochastic nonstationarity is chosen for simplicity.

Using the fact that $E_t[\overline{P}_{t+v+1} - \overline{P}_{t+v}] = (\mu - b)$ for $v = 1, 2, \ldots$, the actual inflation rate in equation (5) can be expressed as

$$\Pi_{t+1} = (\mu - b) + \beta\hat{P}_t, \tag{A3}$$

where $\hat{P}_t$ measures the extent of disequilibrium in the goods market as the difference between the flexible-price level and the actual price level. Note that equation (A3) is identical to equation (13).

Output in the short run is demand determined and can be written as

$$y_t = y_t^d = y_t^s + (y_t^d - y_t^s) = y_t^s + \hat{P}_t. \tag{A4}$$

Note that equation (A4) is identical to equation (10) in that it expresses output as the sum of a trend component and a stationary component that represents the extent of short-run disequilibrium in the goods market. Under our assumption of all shocks to demand and supply being permanent, the stationary component of output, $\hat{P}_t$, is a pure AR(1) process. The stationary component of output is thus unaffected by the characterization of long-run output in terms of a deterministic or a stochastic trend. The only difference in this case, of course, is that detrending output now involves estimating a stochastic rather than a deterministic trend. Once output is appropriately detrended, the inflation rate and the cyclical component of output lagged one period will be perfectly correlated, exactly as in the case of a deterministic trend. Since the cyclical component of output is positively autocorrelated, the rate of inflation will be positively correlated with the cyclical component of output at several lags and leads.

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