The Correlation between Shocks to Output and the Price Level: Evidence from a Multivariate GARCH Model

by

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(abstract)

Previous research indicates that the price-output correlation is time varying. This paper therefore estimates a VAR with a bivariate GARCH error process to obtain quarterly estimates of the price-output correlation for the United States for the period 1876:4-1999:4. The estimated correlation is usually positive before 1945 and zero during 1945-1963. Negative correlations become important only after 1963, but do not become obviously more important than zero correlations. Prior to 1945 the estimated correlation typically is positive during both recessions and expansions. After 1945 the estimated correlation remains largely positive during recessions, but becomes mainly negative during expansions, suggesting that changes in the sign of the price-output correlation primarily are the result of changes in its sign during expansions.

J.E.L. classification: E30, C32

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1. Introduction

In what way has the price level in the United States been related to the level of output? At one time it was generally believed that the correlation between the price level and output is positive and to be useful a macroeconomic model should be able to predict a positive price-output correlation.[1] Hence Mankiw (1989, p. 88) criticizes real business cycle theory because it cannot explain the fact that “inflation tends to rise during booms and fall during recessions.” This general belief in procyclical prices began to change in the early 1990’s after several authors reported finding that detrended measures of the price level and output have a negative correlation during the post-war period. One important early study is that of Kydland and Prescott (1990, p. 17), who conclude that "any theory in which procyclical prices figure crucially in accounting for postwar business cycle fluctuations is doomed to failure. The facts we report indicate that the price level since the Korean War moves countercyclically".[2] This apparent change in the sign of the price-output correlation has resulted in a literature on the cyclical behavior of the price level. Although this literature is largely concerned with measuring the correlation between fluctuations in the price level (or inflation) and fluctuations in output (the price-output correlation), it is also concerned with what determines whether an economy has a positive or negative price-output correlation.

This paper contributes to the literature on the cyclical behavior of prices by presenting time-varying estimates of the price-output correlation. The motivation for this research is the dependence of the estimated price-output correlation on sample period. The difference between pre- and post-war estimates is only one example. There also appears to be some disagreement about when the change in the sign of the correlation occurred. For example, Cooley and Ohanian (1991) report a negative correlation for their entire post-1945 sample; Kydland and Prescott

(1990) argue it has been negative since the end of the Korean War; but Wolf (1991) presents evidence that it did not become negative until after 1973.[3] One advantage of our methodology is that it can estimate when the change in the sign occurred. Another is that it is capable of finding other changes in the sign of the correlation that until now had not been reported in the literature.

Although the previous literature clearly recognizes that the price-output correlation is sample dependent, the typical study implicitly assumes the correlation is fixed over any given sample period. By focusing only on the constant correlation within arbitrarily chosen periods, one may be losing important information about the dynamics of the comovement of output and the price level within and across regimes that might have been misspecified. This paper therefore presents time-varying estimates of the price-output correlation for the United States. This is done by estimating a two-variable VAR in which it is assumed that the disturbances follow a bivariate GARCH process. In the GARCH process, the conditional variance-covariance matrix of the residuals changes over time, allowing quarterly estimates of the contemporaneous price-output correlation coefficient. These estimates allow the identification of periods during which the price-output correlation was generally positive, those during which it was essentially zero, and those during which it was generally negative. It also allows a comparison of whether the price-output correlation is systemically different during periods of recession than it is during periods of recovery.[4]

This paper defines the price-output correlation to be the contemporaneous correlation coefficient between unexpected changes in output and the price level, a definition supported by the work of den Haan (2000). Assuming a fixed residual variance-covariance matrix, he shows that the forecast errors from a VAR can be used to obtain consistent estimates of the price-output correlation as long as the number of lags in the VAR is sufficient to cause the disturbance to be stationary. It does not matter whether the variables in the VAR are stationary and what filter is used to detrend the data.

This paper finds that the price-output correlation is usually positive before 1945, and nearly always close to zero from 1945 through 1963. Beginning in 1963, the frequency of negative correlations increases dramatically. In addition, we find that a zero (or at least statistically insignificant) price-output correlation has been much more common than realized by previous writers. For the entire sample period (1876:4-1999:4), this paper reports a statistically insignificant price-output correlation 57% of the time. For the post-1944 part of the sample, the correlation is zero 70% of the time, significantly positive only 11% of the time, and significantly negative only 19% of the time. The results are supportive of macroeconomic models in which both aggregate demand and aggregate supply shocks can be important in both the short and long runs—that is, models that allow the price-output correlation to be positive, negative, or zero for extended periods of time at forecast horizons up to four years in the future.

Section 2 explains why one should expect the price-output correlation to be time-varying and in doing so offers a brief survey of previous literature on the price-output correlation. Section 3 describes the data and presents the methodology employed by this paper. Section 4 presents and discusses the results, while section 5 offers some conclusions.

2. Reasons for a time-varying price-output correlation

There is some value in asking why one should expect the price-output correlation to be time-varying. The simplest possible explanation for a time-varying price-output correlation is that the relative frequency and sizes of shocks to aggregate supply and aggregate demand change. In a model with flexible prices, during periods when shocks to aggregate supply dominate, the correlation is negative, while during those when shocks to aggregate demand dominate, the correlation is positive.

Several writers, however, confront such a naive connection between the sign of the price-output correlation and the relative importance of supply and demand shocks. Chada and Prasad (1993), Ball and Mankiw (1994) and Judd and Trehan (1995) show that sticky-price models with only demand shocks can yield a negative price-output correlation.[5] These models therefore imply that the price-output correlation could be time-varying because the degree of price stickiness changes over time. Den Haan (2000), however, shows that models with only a demand shock are not capable of generating the pattern of price-output correlations across different forecast horizons which he obtains. Hence his results imply that a changing degree of price stickiness by itself cannot be the only cause of a time-varying price-output correlation.

A more interesting explanation for a time-varying price-output correlation comes from the observation that monetary policy affects the price-output correlation. Pakko (2000) and Cover and Pecorino (2000) present models in which monetary policy can change the sign of the price-output correlation.[6] Since monetary policy is likely to be systematically different during recessions and expansions, their models imply that the price-output correlation is time-varying.

Pakko (2000) obtains his results by simulating a shopping time monetary model with endogenous monetary policy. He finds that the price-output cospectrum is negative at all frequencies in response to a productivity shock if there is a constant growth money rule. But if it is assumed that monetary policy is implemented in a way that makes the money stock procyclical, then Pakko finds that the price-output cospectrum can be made positive at all frequencies. Since the degree to which monetary policy is procyclical changes with economic conditions, Pakko's model implies a time-varying price-output correlation.

Cover and Pecorino (2000) examine an *IS-LM-AS* macroeconomic model with optimal monetary policy. They find that the more successful monetary policy is in offsetting aggregate demand shocks, the less likely it is that the price-output correlation is positive. Whether or not successfully offsetting aggregate demand shocks causes the price-output correlation to be close to zero or to become negative depends upon how the monetary authority responds to supply shocks. The greater the degree to which the monetary authority allows supply shocks to affect output, the more likely it is that the correlation is close to zero. But if the monetary authority tries to prevent *temporary* supply shocks from affecting output, the more likely it is that the price-output correlation is negative. If the degree to which the monetary authority has been successful at offsetting aggregate demand shocks, as well as the degree to which the monetary authority has allowed temporary supply shocks to affect output has changed over time, then Cover and Pecorino's model implies that the price-output correlation is time-varying.

3. Data and Methodology

The US quarterly time series data on real GDP and nominal GDP are from the United States Department of Commerce for the period 1959:1-1999:4 and Balke and Gordon (1986) for the period 1875:1-1959:1. The data sets were spliced together so that the quarterly growth rates of real and nominal GDP through the first quarter of 1959 are equal to those in the Balke-Gordon data, while those beginning with the second quarter of 1959 are equal to those in the Department of Commerce data.[7] The price level is defined to be the GDP deflator.

In order to show that the correlation coefficient is time-varying, we start our analysis with *constant* conditional correlations between output and price over a variety of sample periods, as has been assumed in all previous studies. The constant conditional correlations are calculated from the residuals of an unrestricted VAR(p) process governing quarterly measures of output (y_t) and the price level (P_t):

$$\boldsymbol{X}_t = \boldsymbol{B}_0 + \sum_{j=1}^{p} \boldsymbol{B}_j \boldsymbol{X}_{t-j} + \boldsymbol{e}_t, \tag{1}$$

where $\boldsymbol{X}_t = [\, y_t \;\; P_t \,]'$ and the order of the VAR, p, is set to be the minimum lag length that renders the residual vector $\hat{\boldsymbol{e}}_t$ to be serially uncorrelated and stationary. Equation (1) is estimated by OLS. The Ljung-Box Q-test is used to test for up to 12 lags of serial correlation in $\hat{\boldsymbol{e}}_t$ and the Dickey-Fuller test is used to test the stationarity of $\hat{\boldsymbol{e}}_t$. Because the lag length employed (p) is sufficiently long to eliminate serial correlation, the particular correlation coefficient we are examining is that between changes in output and the price level that cannot be explained by the past behavior of either variable. Hence the correlation coefficient examined here is similar to the one examined by den Haan (2000) and is a consistent estimate of the output-price correlation.

Previous studies [e.g., Kydland and Prescott (1990), Wolf (1991), Cooley and Ohanian (1991), Smith (1992), Backus and Kehoe (1992), and Chada and Prasad (1993)] have tried different measures of variables to calculate the price-output correlation, and different filtering methods to obtain stationarity for the variables. The advantage of using the VAR prediction errors to examine price-output correlations is that, as proved by den Haan (2000), it does not require assumptions about the order of integration or the types of assumptions needed for VAR decompositions, as long as (1) has sufficient lags so that $\boldsymbol{e}_t$ is not integrated. Nevertheless, as den Haan points out, if the variables in the VAR are correctly filtered, then the resulting estimates of the price-output correlation are more efficient. Hence, for cautiousness and to check the robustness of our results, we estimated VAR's using several filters commonly seen in the literature.[8] Table 1 reports the results from two specifications. The "Least Restricted Model" uses the logarithms of output and the price level and includes 5 lags. The "Best Model" is selected by the augmented Dickey-Fuller and KPSS tests. The results of these tests, presented in Appendix A, show that output growth is better modeled as level stationary and inflation is better modeled as trend stationary. Therefore, the "Best Model" uses demeaned and detrended output growth and inflation and includes 4 lags. The results for the other filtered data sets are very similar to those presented in Table 1 and do not qualitatively change any of our conclusions.[9]

The first row of Table 1 shows that the estimated correlation coefficient for the entire sample period is positive and is approximately equal to 0.21. Following Friedman and Schwartz (1982), we further divide the sample into four sub-periods. Prior to 1901 the estimated correlation coefficient is positive but small enough to be essentially zero; during the period 1901-1928 it increases to approximately 0.31; during the period 1929-1946 it increases further to approximately 0.41; and finally during the post-war period it falls below 0.05, once again essentially zero. If [following Cooley and Ohanian (1991)] we look at two sub-periods after WWII, we find that the correlation coefficient is negative (approximately -0.08) for the period 1954:1-1973:1, but essentially zero for the period 1966:1-1999:4.

The results reported in Table 1, as do the results of the previous literature (summarized in Appendix B), clearly show that the price-output correlation is time variant and therefore lend support to the idea that the price-output correlation should be estimated in a manner that allows it to be time-varying. The most widely used time-varying covariance model is the multivariate GARCH model. Several parsimonious specifications of the multivariate GARCH model are available [see Kroner and Ng (1998)]. This paper employs a specification originally proposed by Baba, Engle, Kraft, and Kroner, called the BEKK model [see Engle and Kroner (1995)]. The important feature of this specification is that it builds in sufficient generality and at the same time, it does not require estimation of many parameters. More importantly, the BEKK model guarantees by construction that the variance-covariance matrix in the system is positive definite.

Specifically, assume that the disturbance, $\boldsymbol{e}_t$, in equation (1) is a conditionally zero-mean Gaussian process with time-varying conditional variance-covariance matrix $\boldsymbol{H}_t$: $\boldsymbol{e}_t \mid \boldsymbol{\Omega}_{t-1} \sim N(0, \boldsymbol{H}_t)$, where $\boldsymbol{\Omega}_{t-1}$ is the information set at time $t-1$. Consider the following GARCH (1,1) process,

$$\boldsymbol{H}_t = \boldsymbol{\Gamma}'\boldsymbol{\Gamma} + \boldsymbol{G}'\boldsymbol{H}_{t-1}\boldsymbol{G} + \boldsymbol{A}'\boldsymbol{e}_{t-1}\boldsymbol{e}_{t-1}'\boldsymbol{A}, \qquad (2)$$

where the constant matrix $\boldsymbol{\Gamma}$ is restricted to be upper triangular so that the terms on the right-hand-side of (2) are all in quadratic form, which guarantees a positive definite $\boldsymbol{H}_t$.

In order to restrict the specification for $\boldsymbol{H}_t$ further to obtain a numerically tractable formulation, we set the constant matrices $\boldsymbol{G}$ and $\boldsymbol{A}$ to be diagonal. These restrictions assume that the conditional covariance (the off-diagonal terms in $\boldsymbol{H}_t$) depends only on the past covariances and not on the variances. Finally, in order to obtain estimates that made economic sense, it was necessary to restrict the diagonal elements of $\boldsymbol{G}$ so that their product would always be positive.[10]

After equations (1) and (2) have been estimated jointly by maximum likelihood estimation (MLE), it is straightforward to obtain a time series of estimated correlation coefficients from the series of $\hat{\boldsymbol{H}}_t$, the estimated $\boldsymbol{H}_t$ obtained from the MLE. For inference, we obtain confidence intervals by a Monte Carlo simulation. This was done in the following way. Consider the following alternative expression of the GARCH process:

$$\boldsymbol{e}_t = \boldsymbol{h}_t'\boldsymbol{v}_t, \qquad (3)$$

where $\boldsymbol{v}_t \sim N(\boldsymbol{0},\boldsymbol{I})$ and $\boldsymbol{h}_t$ is an upper triangular matrix obtained from the Cholesky decomposition of $\boldsymbol{H}_t$ such that $\boldsymbol{h}_t'\boldsymbol{h}_t = \boldsymbol{H}_t$. Therefore, $\boldsymbol{e}_t \mid \boldsymbol{\Omega}_{t-1} \sim N(\boldsymbol{0}, \boldsymbol{H}_t)$, where $\boldsymbol{H}_t$ follows the GARCH process (2). Let $\hat{\boldsymbol{h}}_t$ be the Cholesky decomposition of $\hat{\boldsymbol{H}}_t$. One thousand sets of random matrices $\boldsymbol{v}_{it}$ are drawn from the bivariate normal distribution $N(\boldsymbol{0},\boldsymbol{I})$, i = 1, …, 1000. The generated residuals $\hat{\boldsymbol{e}}_{it} = \hat{\boldsymbol{h}}_t \boldsymbol{v}_{it}$ and the estimated coefficients in (1) are then used to generate the artificial data $\hat{\boldsymbol{X}}_{it}$. Finally, equations (1)-(2) are estimated by the MLE using the artificial data. The resulting $\hat{\boldsymbol{H}}_{it}$, i = 1, …, 1000, from these 1000 simulations are used to construct the one standard deviation band for the original $\hat{\boldsymbol{H}}_t$, which in turn includes the one-standard deviation band for the estimated correlation coefficients.

4. Results

4.1 Results for price-output correlation from one-step-ahead forecast errors.

The series of correlation coefficients obtained directly from estimating equations (1) and (2) are estimates of the time series of correlation coefficients for the stochastic error term in equation (1). Because the residuals obtained from estimating equation (1) are one-step-ahead forecast errors, the results discussed in this subsection could be called results for one-step-ahead forecast errors, but here they are simply referred to as the correlation coefficient. Section 4.3 discusses results for other forecast horizons.

Figure 1 presents the estimates of the time series of correlation coefficients along with the one-standard-deviation band obtained from the simulations obtained using the "best model".[11] As can be seen from Figure 1 there are many periods during which the correlation coefficient is not significantly different from zero. In order to make it easy to identify periods during which the correlation coefficient is either positive or negative, Figure 2 presents a graph in which statistically insignificant values of the correlation coefficient are set to zero. A line segment connects consecutive observations of the estimated correlation only if both observations are significantly different from zero.

Figure 2 shows that the correlation coefficient was largely positive during four distinct periods: (1) the early 1880's; (2) the long period from 1907 through 1944; (3) 1958-59; and (4) 1981-1985. One also can see that there are five periods during which the correlation coefficient was largely negative: (1) 1877:4-1878:3; (2) 1892-1893; (3) 1945:4-1946:3; (4) 1964:1-1978:2, and (5) the last seven years of the sample, 1993:1-1999:4. The estimated price-output correlation coefficient is not significantly different from zero during four relatively long periods. These are the periods 1883-91, 1894-1906, 1947-1957, and 1986-1992. Additionally the correlation is not significantly different from zero during the following shorter periods: 1928-1929, 1959-1963, 1967:4-1969:1 and 1971:2-1974:3.

Table 2 provides a summary of the information in Figure 2 by presenting for various sample periods the frequency of significantly positive, not significantly different from zero, and significantly negative correlations. For the entire sample (1876:4-1999:4), the estimated correlation coefficient is not significantly different from zero over one-half of the time, significantly positive about one-third of the time, and significantly negative only 12% of the time. But the majority of positive observations occur prior to 1945. Hence for the 1945:1-1999:4 period, the correlation is significantly positive only 11% of the time using the best model and 14% of the time using the least restrictive model. It is notable that Table 2 implies that during the post-war period the price-output correlation has not been largely negative, rather it has been largely not significantly different from zero since 1945. Only about 20% of quarters yield a significantly negative correlation, while 70% of quarters have a correlation coefficient not significantly different from zero.

Table 2 clearly shows that negative correlations also are largely a post-1963 phenomenon. During the period 1945:1-1963:4, the correlation was significantly negative only

5% of the time, but significantly negative over one-fourth of the time after 1963. Table 2 suggests that rather than asking why the correlation changed from being positive before the war to being negative after the war, macroeconomists should be asking why it was nearly always zero during 1945-1963, and why negative correlations rarely appeared before 1964.

One possible criticism of the conclusion that the price-output correlation is usually zero is that it is based upon simulated standard deviations that are relatively large. Nearly all of the standard deviations are greater than 0.15, about 90% are greater than 0.17, and over one-half of them are greater than 0.2. This means that the absolute value of an estimated correlation coefficient must be greater than 0.17 to be considered significant. What sort of results do we obtain if we use more modest criteria to determine whether an estimated correlation coefficient is economically significant?

The more modest criteria used here are as follows. (1) A period with a *clearly positive* correlation is one in which the estimated price-output correlation is continuously positive and includes at least four consecutive quarters during which the estimated correlation coefficient is greater than 0.15. (2) A period with a *clearly negative* correlation is one in which the estimated correlation is always negative, but includes at least four consecutive quarters during which the estimated correlation coefficient is less than -0.15. (3) Any period of at least four quarters with an estimated correlation that is continuously less than 0.15 in absolute value, as well as any period that does not meet the above two criteria is considered a period with an *essentially zero* correlation. Tables 3 and 4 present average values of the estimated correlation coefficient for periods with positive and negative price-output correlation coefficients based on these modest criteria.

As shown in Table 3, the above criteria yield 8 periods with a clearly positive price-output correlation. The three longest periods are within the years 1907-1944 and each is longer than 7 years.[12] The fourth longest period with a positive correlation is 1981:1-1985:3, only one quarter shy of being 5 years long, while the fifth longest is the 1879:4-1883:3 period. During each of these five periods, the average value of the price-output correlation is greater than 0.3.

As can be seen from Table 4, the above criteria continue to imply that the price output correlation was rarely negative before 1964. The period with the longest continuously negative price-output correlation (by these criteria) is the 27-quarter period that closes the sample (1993:2-1999:4). For the least restrictive model, the second and third longest periods (1964:1-1967:3 and 1974:4-1978:2) are both one-quarter shy of being four years long. During each of the post-1963 periods with a clearly negative price-output correlation, the average value is between –0.22 and –0.27.

Table 5 presents the shares of quarters with clearly positive, clearly negative, and essentially zero correlations. For the period 1876:4-1944:4, 57% of quarters have a clearly positive correlation, but only 5% have a clearly negative correlation according to the results obtained from the best model. For the sample period 1945:1-1963:4, the share of quarters with a clearly positive correlation declines to 8%, while 87% of quarters have an essentially zero correlation. Finally, for the post-1963 sample, the share of quarters with a negative correlation jumps to 35%, the share with positive correlations increases to 17% , while the share of quarters with an essentially zero correlation falls dramatically to 49% according the results obtained from the best model.

Even though the results in Tables 2 and 5 are somewhat in line with the findings in previous studies with constant correlations, namely positive output-price correlations for the prewar sample and negative for the post-war (or post-1973) period, the methodology used in this paper provides more information about the dynamics of the comovements of output and the price level. For example, we find that negative correlations are largely a post-1963, instead of post-war or post-1973, phenomenon and that the correlation has been insignificantly different from zero much more often than realized by previous studies.

Table 5 raises questions that are similar to those suggested by Table 2. In particular, why did the frequency of positive correlations decline after 1944, why is the correlation almost always zero during 1945-63, and why did negative correlations suddenly become more important after 1963? Like Table 2, Table 5 suggests that macroeconomists should look for models that allow the price-output correlation to be positive, zero, or negative for long periods.

4.2 Comparison of results for recessions and expansions

Since one purpose of macroeconomics is to explain the existence of business cycles, it is natural to ask whether the price-output correlation tends to have different values during recessions and expansions. Since the bars in Figure 2 represent periods of recession, a cursory glance at that figure shows that the correlation is rarely negative and significant during recessions. Table 6, which presents average values of the estimated price-output correlation for periods of recession and recovery, helps make sense of this observation. Using the results for the best model, for the entire sample the average correlation is 0.134 during recessions and only 0.087 during expansions. But these averages mean nothing because of the tendency for the correlation to be positive during pre-1945 expansions and negative during post-1945 expansions. For the pre-1945 sample, the average correlation is 0.150 during recessions and 0.228 during expansions.

The results are somewhat different for the period after 1944. Notice from Figure 1 that during post-war recessions the estimated correlation is typically positive (though not always significant). Figure 2 shows that during the post-war period significantly negative price-output correlations usually appear only during expansions. Indeed, they appear only twice during post-war recessions: during the last two quarters of 1970 and during the fourth quarter of 1974 and the first quarter of 1975. The last two rows of Table 6 show that during the post-war period the correlation coefficient is more likely to be positive during recessions and more likely to be negative during expansions. The average value of the correlation during 1945:1-1963:4 is 0.084 during recessions and –0.036 during expansions. Because the correlation coefficient was essentially zero for most of this period, the small size of these numbers is not surprising. After 1963, however, the average value of the price-output correlation increases to 0.109 during recessions and falls to –0.059 during expansions.

Combining the results in Table 6 with those in Tables 2 and 5 it is reasonable to conjecture that the main reason for changes in the price-output correlation is changes in its value during periods of expansion. Although after 1944 the average value of the correlation during recessions declined, it continues to be mostly positive during periods of recession. On the other hand, the average value of the correlation during expansions not only declined after 1944, but changed from being mostly positive to mostly negative after 1963.

Table 7 drives home this point by presenting for various periods the shares of quarters in which the point estimates of the price-output correlation are positive during recessions and expansions. Notice that for the best model the share of quarters with a positive price-output correlation during recessions varies only from 71% to 74%. But for expansions, the share of quarters with a positive price-output correlation declines from 83% before 1945 to 58% for

1945-1963, and then to 43% for 1964-1999. Both before and after the war more than two-thirds of recessionary quarters have a positive price-output correlation. The results in Table 7 clearly suggest that the tendency of the price-output correlation to be positive during recessions has changed very little, while its tendency to be positive during expansions has declined dramatically since 1945.

4.3 Results for other forecast horizons.

As mentioned above, den Haan (2000) shows that examining the price-output correlation at different forecast horizons can provide information about what types of macroeconomic models are consistent with the data. When using quarterly data for the period 1960:2-1997:2, he finds that the price-output correlation is approximately zero at relatively short forecast horizons and negative at relatively distant forecast horizons.[13] To see if Den Haan's results hold up under our methodology, in this section we present estimates of the price-output correlation at two additional forecast horizons. Appendix C derives and presents the formula for calculating the price-output correlation, which depends on both the VAR coefficients and the parameters in the GARCH process, at various forecast horizons.

Figure 3 presents the 8-step-ahead and 16-step-ahead price-output correlations along with the one-step-ahead price-output correlations previously presented in Figure 1. Figure 3 shows that the 16-step ahead correlation is nearly always less than the 8-step ahead correlation, which in turn is nearly always less than the 1-step ahead correlation. Hence a negative correlation is much more prevalent for both the 8- and 16-step-ahead forecast periods than for the one-step-ahead period. This result is broadly consistent with the findings of den Haan (2000) and Pakko (2000), both of whom find that a positive correlation tends to be a short-run or high-frequency phenomenon, while negative correlations tend to have a longer-run or lower-frequency manifestation. But it is also clear from Figure 3 that the 8- and 16-step-ahead correlations are not *always* less than the one-step-ahead correlation. This is particularly true during the period after 1963 when all three correlations are often approximately equal. Indeed, during much of the post-1963 period, there is very little difference between the three correlations. This implies that often since 1963 the relative importance of demand and supply shocks often has been similar in the long and short runs.

5. Summary and conclusions

This paper finds that the price-output correlation has changed often since 1876. The estimates suggest that the correlation is usually positive before 1945, and nearly always close to zero from 1945 through 1963. After 1963, however, the frequency of negative correlations increases dramatically. Although periods with a consistently positive or negative price-output correlation do not coincide in any obvious way with the business cycle, the paper does find that after 1944 the price-output correlation is rarely negative during recessions and rarely positive and significant during expansions. This suggests that it is reasonable to conjecture that the price-output correlation behaves differently after 1944 mainly because it is behaving differently during periods of expansion.

These apparent changes in the price-output correlation coefficient suggest that either the relative importance of demand-side and supply-side shocks has changed over time, the reaction of policy makers to such shocks has changed, or both. The results also suggest that economists have overemphasized the idea that there has been a negative correlation between prices and output during the post-war period. Macroeconomists should not look for models that limit the sign or size of the price-output correlation.

References

Backus, David K. and Patrick J. Kehoe, "International Evidence on the Historical Properties of Business Cycles." *American Economic Review* 82, September 1992: pp. 864-888.

Balke, Nathan S., and Robert J. Gordon, "Appendix D: Historical Data," *The American Business Cycle: Continuity and Change* edited by Robert J. Gordon. Chicago, University of Chicago Press: 1986, pp. 781-850.

Ball, Lawrence and N. Gregory Mankiw, "A Sticky-Price Manifesto." *Carnegie-Rochester Conference Series on Public Policy*, Volume 41, December 1994, pp. 127-151.

Blanchard, Olivier J. and Danny Quah, "The Dynamic Effects of Aggregate Demand and Supply Disturbances." *American Economic Review* 79*, September 1989, pp. 655-73.*

Blanchard, Olivier J. and Stanley Fishcer, *Lectures on Macroeconomics*, MIT Press, Cambridge, MA: 1989.

Burns, Arthur and Wesley C. Mitchell, *Measuring Business Cycles,* National Bureau of Economic Research, 1946.

Chada, Bankim and Eswar Prasad, "Interpreting the Cyclical Behavior of Prices." *IMF Staff Papers* 40, June 1993: pp. 266-298.

Cooley, T. F. and L. H. Ohanian, "The Cyclical Behavior of Prices" *Journal of Monetary Economics* 28, August 1991: pp. 25-60.

Cover, James P. and Paul Pecorino, "Optimal Monetary Policy and the Correlation between Prices and Output." University of Alabama Department of Economics, Finance and Legal Studies Working Paper # WP00-08-05, August 2000.

den Haan, Wouter J., "The Comovement between Output and Prices." *Journal of Monetary Economics* 46, August 2000: pp. 3-30.

Dickey, David A. and Sastry G. Pantula, “Determining the Order of Differencing in Autoregressive Processes.” *Journal of Business and Economic Statistics* 5(4), 1987: 455-461.

Engle, Robert F. and Kenneth F. Kroner. “Multivariate Simultaneous Generalized ARCH.” *Econometric Theory* 11, 1995, pp. 122-150.

Flodén, Martin. "Endogenous Monetary Policy and the Business Cycle." *European Economic Review* 44, August 2000: 1409-29.

Friedman, Milton, “Nobel Lecture: Inflation and Unemployment.” *Journal of Political Economy* 85, No. 3. (Jun., 1977), pp. 451-472.

Friedman, Milton and Anna J. Schwartz, *Monetary Trends in the United States and the United Kingdom: Their Relation to Income, Prices and Interest Rates, 1867-1975*. University of Chicago Press, Chicago: 1982.

Gavin, William T., and Finn E. Kydland, "Endogenous Money Supply and the Business Cycle." *Review of Economic Dynamics* 2, April 1999, pp. 347-369.

Hartley, James E., “Real Myths and a Monetary Fact,” *Applied Economics* 31, November 1999, pages 1325-1329.

Hansen, Alvin H., *Business Cycles and National Income.* W. W. Norton, New York: 1951.

Hobijn, Bart, Philip Hans Franses, and Marius Ooms, “Generalization of the KPSS-test for Stationarity.” *Econometric Institute Report*, No. 9802/A, 1998.

Judd, John. P. and Bharat Trehan, "The Cyclical Behavior of Prices: Interpreting the Evidence." *Journal of Money, Credit and Banking* 27, August 1995: pp. 789-797.

Kroner, Kenneth F. and Victor K. Ng, “Modeling Asymmetric Comovements of Asset Returns.” *The Review of Financial Studies* Winter 1998, Vol. 11, No. 4, pp. 817-844.

Kydland, Finn E. and Edward C. Prescott, "Business Cycles: Real Facts and a Monetary Myth." *Federal Reserve Bank of Minneapolis Quarterly Review* 14 (Spring) 1990: pp. 3-18.

Mankiw, N. Gregory, "Real Business Cycles: A New Keynesian Perspective." *Journal of Economic Perspectives* 3, Summer 1989: pp. 79-90.

McConnell, Margaret M. and Gabriel Perez-Quiros, “Output Fluctuations in the United States: What Has Changed Since the Early 1980’s?” *American Economic Review* 90, December 2000, pp. 1464-1476.

Ng, Serena and Pierre Perron, “Unit Root Tests in ARMA Models with Data-Dependent Methods for the Selection of the Truncation Lag.” *Journal of the American Statistical Association* Vol. 90, No. 429, 1995: 268-280.

Pakko, Michael R., "The Cyclical Relationship between Output and Prices: An Analysis in the Frequency Domain." *Journal of Money, Credit and Banking* 32, August 2000, Part 1, pp. 382-399.

Smith, R. Todd, "The Cyclical Behavior of Prices." *Journal of Money, Credit, and Banking* 24, November 1992: pp. 413-30.

Spencer, David E., "Interpreting the Cyclical Behavior of the Price level in the U.S." *Southern Economic Journal* 63, July 1996: pp. 95-105.

Wolf, Holger C. "Procyclical Prices: A Demi-Myth?" *Federal Reserve Bank of Minneapolis Quarterly Review* 15 (Spring) 1991: pp. 25-28.

Footnotes

1. See, for example, Burns and Mitchell (1946: p.101), Hansen (1951, pp. 4-5), and Blanchard and Fischer (1989: p. 20). It should be noted that before 1990 not all economists thought that prices are always procyclical. The most important exception is Milton Friedman (1977) who in his Nobel Prize lecture explores the possibility of countercyclical prices. See also Friedman and Schwartz (1982: p. 402), who report a negative correlation between output and the price level for both the United States and the United Kingdom

2 . Other early contributors to this literature are Backus and Kehoe (1992), Cooley and Ohanian (1991), Smith (1992), and Wolf (1991).

3. There is also ample evidence that the signs and sizes of other correlations depend on sample period. See, for example, Hartley (1999).

4. It is worth noting that by allowing the residuals to follow a GARCH process, the empirical model employed here implicitly takes into account changes in the variance of output, such as that discussed by McConnell and Perez-Quiros (2000).

5. Spencer (1996) uses similar arguments to show that a negative price-output correlation is consistent within an AD/AS model in which demand shocks affect the price level permanently and output temporarily, while supply shocks have permanent effects on both output and the price level.

6. Flodén (2000) and Gavin and Kydland (1999) also present models in which monetary policy affects the price-output correlation.

7. This is equivalent to multiplying the logarithms of the Ballke-Gordon data by a constant so that the Balke-Gordon observation for 1959:1 equals that for the Department of Commerce Data. Michael Pakko provided the authors with the Balke-Gordon data.

8. The complete set of filters considered were (1) demeaning and detrending, (2) first-differencing, (3) demeaning and detrending the first-differenced data, (4) using first-differenced output and the second differenced price level, and (5) demeaning and detrending first-differenced output and second differenced price level. As is stated in the text, only the third specification was not rejected by the data. The results for the filtered data are too similar to those reported in the text to be worth reporting here. We chose to report only the results for the third specification (the "Best Model") and the unfiltered data ("least restrictive model") because each of the other filters is implicitly imposing a constraint on the data rejected by our specification tests.

9. In addition to the filters mentioned above in footnote 8, we also filtered the data with the Hodrick-Prescott (H-P) filter. Again, the results obtained using this filter are very similar to those reported in the text.

10. If the product of the diagonal elements in ***G*** is allowed to be negative, it tends to dominate the other terms in equation (2), causing the covariance term in (2) to be negatively related to its lagged value. Hence any quarter with a positive covariance is almost always followed by one with a negative covariance, which in turn is followed by one with a positive covariance. The authors do not believe such a pattern makes economic sense. Furthermore, estimates in which the diagonal elements of ***G*** were not restricted to have a positive product often did not converge.

11. The plot of the correlation coefficients obtained from the least restrictive model (that is, using only the logarithms of the data) is very similar to Figure 1. As is made clear from Tables 2-6, the best model is slightly more likely to yield a price-output correlation that is insignificantly different from zero.

12. The period 1907:1-1914:3 is separated from the period 1915:1-1917:4 in Table 3 because the estimated correlation for 1914:4 is –0.154 using the least restrictive model and -0.213 using the best model.

13. With monthly data den Haan finds that correlations are positive at short forecast horizons and gradually become negative as the horizon lengthens.

Appendix A

We identify the order of integration for each variable by using the augmented Dickey-Fuller (ADF) tests, which test the null hypothesis of a unit root, and the generalized KPSS test [Hobijn, Franses, and Ooms (1998)], which tests the null hypothesis of stationarity against the alternatives of nonstationarity. Following a suggestion by Dickey and Pantula (1987), the unit root tests are first conducted with two roots, and if two roots are rejected, then single unit root is tested. Column (1) of Table A1 shows that the ADF test rejects the null of a unit root for both second-differenced series and the KPSS test fails to reject the null of zero-mean stationary for both series. That is, the second-differenced series are better modeled as zero-mean stationary processes. Column (2) shows that the ADF test rejects the null of a unit root for both first-differenced series. The KPSS test fails to reject the null of level stationary for output growth and the null of trend stationary for inflation. That is, output growth is better modeled as a stationary process with a drift, while inflation is better modeled as a stationary with a drift and a deterministic time trend. In column (3), the results show that for the log-price level, the ADF test fails to reject the null of unit root and the KPSS tests reject the null of stationarity, i.e., the log price is better modeled as I(1). For the log-output level, the ADF test rejects the null of unit root at the 5% level but not at the 1% level. However, based on the facts that the KPSS tests reject the null of stationarity and that the estimated coefficient on the first lag is 0.968, we model the log-output as an I(1) series.

Table A1
Unit Root Tests

	(1)		(2)		(3)		
	Differenced Output growth	Differenced Inflation	Output growth	Inflation	Log Output	Log Price	5% Critical Value
ADF [a]	-13.747	-12.518	-10.423	-8.400	-3.629	-2.394	-3.432
KPSS (Null: zero-mean stationary)	0.047	0.090	11.825	2.618	360.742	60.156	1.656
KPSS (Null: level stationary)	-------	-------	0.042	0.692	3.188	2.966	0.460
KPSS (Null: trend stationary)	-------	-------	-------	0.075	0.174	0.588	0.148

a. This is the ADF t-test. The ADF ρ-test yields the same conclusion. The lags in the ADF test are selected by the AIC criterion. We also use Schwartz's criterion and a sequence of 5% t tests for the significance of coefficients on additional lags, as suggested by Ng and Perron (1995). The results with these two criteria do not change the conclusions here.

Appendix B

Table B1
Results of Previous Writers.

Author	Filter	Sample	ρ^a	Main Conclusion
Kydland and Prescott (1990)	Hodrick-Prescott	1954-1989, quarterly	-0.55	Negative correlation since Korean War.
Wolf (1991)	Hodrick-Prescott	1957-1989, quarterly	0.09	Positive correlation before 1973, negative after 1973.
		1973-1989, quarterly	-0.40	
Cooley and Ohanian (1991)	Hodrick-Prescott	Annual: 1870-1900	0.24	The only consistent positive correlation is between two world wars.
		1900-1928	0.24	
		1928-1946	0.77	
		1949-1975	-0.58	
		Quarterly: 1948:2-1987:2	-0.57	
		1954:1-1973:1	-0.36	
		1966:1-1987:2	-0.68	
	log-differencing	Annual: 1870-1900	0.05	
		1900-1928	0.12	
		1928-1946	0.67	
		1949-1975	-0.07	
		Quarterly: 1948:2-1987:2	-0.06	
		1954:1-1973:1	-0.05	
		1966:1-1987:2	-0.23	
	linear detrending	Annual: 1870-1900	-0.17	
		1900-1928	-0.12	
		1928-1946	0.73	
		1949-1975	-0.53	
		Quarterly: 1948:2-1987:2	-0.67	
		1954:1-1973:1	-0.69	
		1966:1-1987:2	-0.87	

a. Estimate of contemporaneous correlation coefficient.

Table B1 - continued.

Author	Filter	Sample	ρ^a	Main Conclusion
Cooley and Ohanian (1991)	First-differenced output & second-differenced price	Annual: 1871-1975	-0.03	
		1871-1910	-0.13	
		1928-1946	0.28	
		Quarterly: 1948:2-1987:1	0.03	
		1954:1-1973:1	-0.03	
		1966:1-1987:1	0.05	
Smith (1991)	Hodrick-Prescott	Annual: 1869-1909	0.23	Positive correlation from the late 19th century until the WWII, except period around WWI. Negative correlation for the post-Depression period, except a period in the 1950's or 1960's.
		1910-1929	0.03	
		1930-1945	0.49	
		1946-1983	-0.68	
	First differencing	Annual: 1869-1909	0.04	
		1910-1929	-0.02	
		1930-1945	0.84	
		1946-1983	-0.54	
Backus and Kehoe (1992)	Hodrick-Prescott	Annual: Prewar	0.22	
		Interwar	0.72	
		Postwar	-0.30	
	Log differencing	Prewar	0.13	
		Interwar	0.37	
		Postwar	-0.25	

a. Estimate of contemporaneous correlation coefficient.

Table B1 – continued

Author	Filter	Sample	ρ^a	Main Conclusion
Chada and Prasad (1993)	Log differencing	1947-1989, quarterly.	-0.07	Similarly filtered price and output have negative correlations. Inflation rate has positive correlation with output using other filters.
	Linear detrending		-0.69	
	Linear detrending with 1973 break		-0.10	
	Hodrick-Prescott		-0.19	
	Inflation & trended output		0.21	
	Inflation & trended output with 1973 break		0.13	
	Inflation & Hodrick-Prescott filtered output		0.16	
	Inflation & Beveridge-Nelson stationary component of output		From 0.01 to 0.20 for different ARMA models	
	Inflation & Blanchard-Quah stationary component of output		0.67	
Den Haan (2000)	Forecast errors from VAR	1948-1997, quarterly & monthly.	Up to 0.4 in short forecasting horizons and down to –0.6 in long forecasting horizons.	Correlation is approximately zero at relatively short forecast horizons and negative at relatively distant forecast horizons.
Pakko (2000)	Cospectrum of Real GNP and GNP Deflator after applying H-P filter and first-differencing to Gordon-Balke data set.	Quarterly 1875:1-1914:4	Positive at low, near zero at high frequencies	Correlation changes sign after the War because of changes in co-spectrum at low frequencies.
		1920:1-1940:4	Positive at all frequencies	
		1950:1-1994:4	Negative at low frequencies	

a. Estimate of contemporaneous correlation coefficient.

Appendix C

To find the time-varying conditional forecast errors, consider the Wald representation of equation (1):

$$\boldsymbol{X_t} = \alpha_0 + \sum_{j=1}^{\infty} \alpha_j \boldsymbol{e}_{\boldsymbol{t}+1-j}, \qquad \alpha_1 = \boldsymbol{I}.$$

Let the s-step-ahead forecast error of $\boldsymbol{X_t}$ be $\boldsymbol{F_t^s}$. Then

$$\boldsymbol{F_t^s} = \boldsymbol{X_t} - \boldsymbol{E}(\boldsymbol{X_t} \mid t-s) = \boldsymbol{e_t} + \alpha_2 \boldsymbol{e}_{\boldsymbol{t}-1} + \alpha_3 \boldsymbol{e}_{\boldsymbol{t}-2} + \ldots + \alpha_{\boldsymbol{s}} \boldsymbol{e}_{\boldsymbol{t}-\boldsymbol{s}+1},$$

and $\boldsymbol{E}(\boldsymbol{F_t^s} \mid t-s) = 0$. The conditional variance-covariance matrix of $\boldsymbol{F_t^s}$ is:

$$\begin{aligned}
&\boldsymbol{V}(\boldsymbol{F_t^s} \mid t-s) \\
&= \boldsymbol{E}[(\boldsymbol{e_t} + \alpha_2 \boldsymbol{e}_{\boldsymbol{t}-1} + \alpha_3 \boldsymbol{e}_{\boldsymbol{t}-2} + \ldots + \alpha_{\boldsymbol{s}} \boldsymbol{e}_{\boldsymbol{t}-\boldsymbol{s}+1})(\boldsymbol{e_t} + \alpha_2 \boldsymbol{e}_{\boldsymbol{t}-1} + \alpha_3 \boldsymbol{e}_{\boldsymbol{t}-2} + \ldots + \alpha_{\boldsymbol{s}} \boldsymbol{e}_{\boldsymbol{t}-\boldsymbol{s}+1})' \mid t-s] \\
&= \boldsymbol{E}[(\boldsymbol{e_t}\boldsymbol{e_t'} + \alpha_2 \boldsymbol{e}_{\boldsymbol{t}-1}\boldsymbol{e}'_{\boldsymbol{t}-1}\alpha_2' + \ldots + \alpha_{\boldsymbol{s}} \boldsymbol{e}_{\boldsymbol{t}-\boldsymbol{s}+1}\boldsymbol{e}'_{\boldsymbol{t}-\boldsymbol{s}+1}\alpha_{\boldsymbol{s}}') \mid t-s],
\end{aligned}$$

where the second equality uses the fact that $\boldsymbol{e_t}$ is serially uncorrelated. Let $\boldsymbol{e_t}\boldsymbol{e_t'} = \boldsymbol{H}_t + \boldsymbol{w}_t$, where $\boldsymbol{w}_t$ is a white noise process. Then

$$\begin{aligned}
&\boldsymbol{V}(\boldsymbol{F_t^s} \mid t-s) \\
&= \boldsymbol{E}[(\boldsymbol{H_t} + \boldsymbol{w_t}) + \alpha_2(\boldsymbol{H}_{\boldsymbol{t}-1} + \boldsymbol{w}_{\boldsymbol{t}-1})\alpha_2' + \ldots + \alpha_{\boldsymbol{s}}(\boldsymbol{H}_{\boldsymbol{t}-\boldsymbol{s}+1} + \boldsymbol{w}_{\boldsymbol{t}-\boldsymbol{s}+1})\alpha_{\boldsymbol{s}}' \mid t-s] \\
&= \boldsymbol{E}[\boldsymbol{H_t} + \alpha_2 \boldsymbol{H}_{\boldsymbol{t}-1}\alpha_2' + \ldots + \alpha_{\boldsymbol{s}} \boldsymbol{H}_{\boldsymbol{t}-\boldsymbol{s}+1}\alpha_{\boldsymbol{s}}' \mid t-s] \\
&= \boldsymbol{E}(\boldsymbol{H_t} \mid t-s) + \alpha_2 \boldsymbol{L} \cdot \boldsymbol{E}(\boldsymbol{H_t} \mid t-s+1)\alpha_2' + \alpha_3 \boldsymbol{L}^2 \cdot \boldsymbol{E}(\boldsymbol{H_t} \mid t-s+2)\alpha_3' + \ldots + \alpha_s \boldsymbol{L}^{s-1}\boldsymbol{E}(\boldsymbol{H_t} \mid t-1)\alpha_s',
\end{aligned}$$

where $\boldsymbol{L}$ is the lag operator. Taking the conditional expectation of equation (2), we have

$$\boldsymbol{E}(\boldsymbol{H_t} \mid t-k) = \varGamma'\varGamma + \boldsymbol{G'L} \cdot \boldsymbol{E}(\boldsymbol{H_t} \mid t-k+1)\boldsymbol{G} + \boldsymbol{A'L} \cdot \boldsymbol{E}(\boldsymbol{H_t} \mid t-k+1)\boldsymbol{A}.$$

Using the fact that $\boldsymbol{E}(\boldsymbol{H_t} \mid t-1) = \boldsymbol{H_t}$ and solving the above equation recursively for $k = 2, \ldots, s$, $\boldsymbol{V}(\boldsymbol{F_t^s} \mid t-s)$ can be derived from $\boldsymbol{H_t}$.

Table 1:

Constant conditional correlation cofficients between the logarithms of output and the price level

Sample period	Best Model	Least Restrictive Model
1876:4-1999:4	0.213	0.218
1876:4 1900:4	0.056	0.065
1901:1-1928:4	0.306	0.307
1929:1-1946:4	0.407	0.414
1948:2-1999:4	0.046	0.032
1954:1-1973:1	-0.077	-0.076
1966:1-1999:4	0.004	-0.003

Table 2

Share of quarters when correlation coefficient is positive, zero, and negative

Sample period	Share of quarters significantly positive		Share of quarters not significantly different from zero		Share of quarters significantly negative	
	Best Model	Least Restrictive Model	Best Model	Least Restrictive Model	Best Model	Least Restrictive Model
1876:4-1999:4	31%	34%	57%	54%	12%	12%
1876:4-1944:4	47%	51%	47%	44%	6%	5%
1945:1-1999:4	11%	14%	70%	65%	19%	21%
1945:1-1963:4	8%	9%	87%	86%	5%	5%
1964:1-1999:4	14%	16%	61%	55%	26%	29%

Table 3

Average value of estimated price-output correlation coefficient for various sample periods during which the correlation is clearly positive[a]

Sample period	Average Estimated Correlation		Length of period
	Best Model	Least Restrictive Model	
(1) 1879:4 to 1883:3	0.352	0.354	16
(2) 1893:4 to 1895:2	0.236	0.217	7
(3) 1907:1 to 1914:3	0.317	0.333	31
(4) 1915:1 to 1927:4	0.328	0.319	52
(5)1930:1 to 1944:3	0.384	0.443	59
(6) 1958:1 to 1959:2	0.295	0.269	6
(7) 1978:3 to 1980:2	0.203	0.203	12
(8) 1981:1 to 1985:3	0.323	0.358	19

a. A clearly positive period is one in which the correlation coefficient is continuously positive and includes at least four consecutive observations greater than 0.15.

Table 4

Average value of estimated price-output correlation coefficient for various sample periods in which the correlation is clearly negative[a]

Sample period	Average Estimated Correlation		Length of period
	Best Model	Least Restrictive Model	
(1) 1877:4 to 1878:3	-0.295	-0.257	4
(2) 1891:4 to 1893:2	-0.388	-0.377	7
(3) 1903:1 to 1904:1	-0.190	-0.203	4
(4) 1945:4 to 1946:3	-0.437	-0.423	4
(5a) 1964:1 to 1965:1	-0.235	___	5
(5b) 1964:1 to 1967:3	___	-0.223	15
(6) 1969:2 to 1971:1	-0.209	-0.246	8
(7) 1974:4 to 1978:2	-0.267	-0.267	15
(8) 1993:2 to 1999:4	-0.244	-0.221	27

a. A clearly negative period is one in which the correlation coefficient is continuously negative and includes at least four consecutive observations less than -0.15.

Table 5

Share of observations when correlation is clearly positive, essentially zero and clearly negative[a]

Sample Period	Share of quarters clearly positive		Share of quarters essentially zero		Share of quarters clearly negative	
	Best Model	Least Restrictive Model	Best Model	Least Restrictive Model	Best Model	Least Restrictive Model
1876:4-1999:4	38%	40%	48%	43%	14%	17%
1876:4-1944:4	57%	60%	38%	34%	5%	5%
1945:1-1999:4	14%	15%	62%	54%	24%	31%
1945:1-1963:4	8%	8%	87%	87%	5%	5%
1964:1-1999:4	17%	19%	49%	37%	35%	44%

a. A clearly positive period is one in which the correlation coefficient is continuously positive and includes at least four consecutive observations greater than 0.15. A clearly negative period is defined analogously. All other periods have an essentially zero correlation.

Table 6

Average values of price-output correlation during recessions and expansions

Sample	Recession		Expansion	
	Best Model	Least Restrictive Model	Best Model	Least Restrictive Model
1876:4-1999:4	0.134	0.149	0.087	0.097
1876:4-1944:4	0.150	0.167	0.228	0.250
1945:1-1963:4	0.084	0.072	-0.036	-0.016
1964:1-1999:4	0.109	0.104	-0.059	-0.045

Table 7

Share of quarters with a positive price-output correlation during recessions and expansions.[a]

Sample	Recession		Expansion	
	Best Model	Least Restrictive Model	Best Model	Least Restrictive Model
1876:4-1999:4	74%	77%	64%	63%
1876:4-1944:4	74%	78%	83%	84%
1945:1-1963:4	73%	80%	58%	52%
1964:1-1999:4	71%	70%	43%	43%

a. In this table a quarter is defined as having a positive price-output correlation if its point estimate is positive.

Figure 1
Estimated Values of Rho with One Standard Deviation Bounds
(Shaded areas represent NBER periods of recession.)

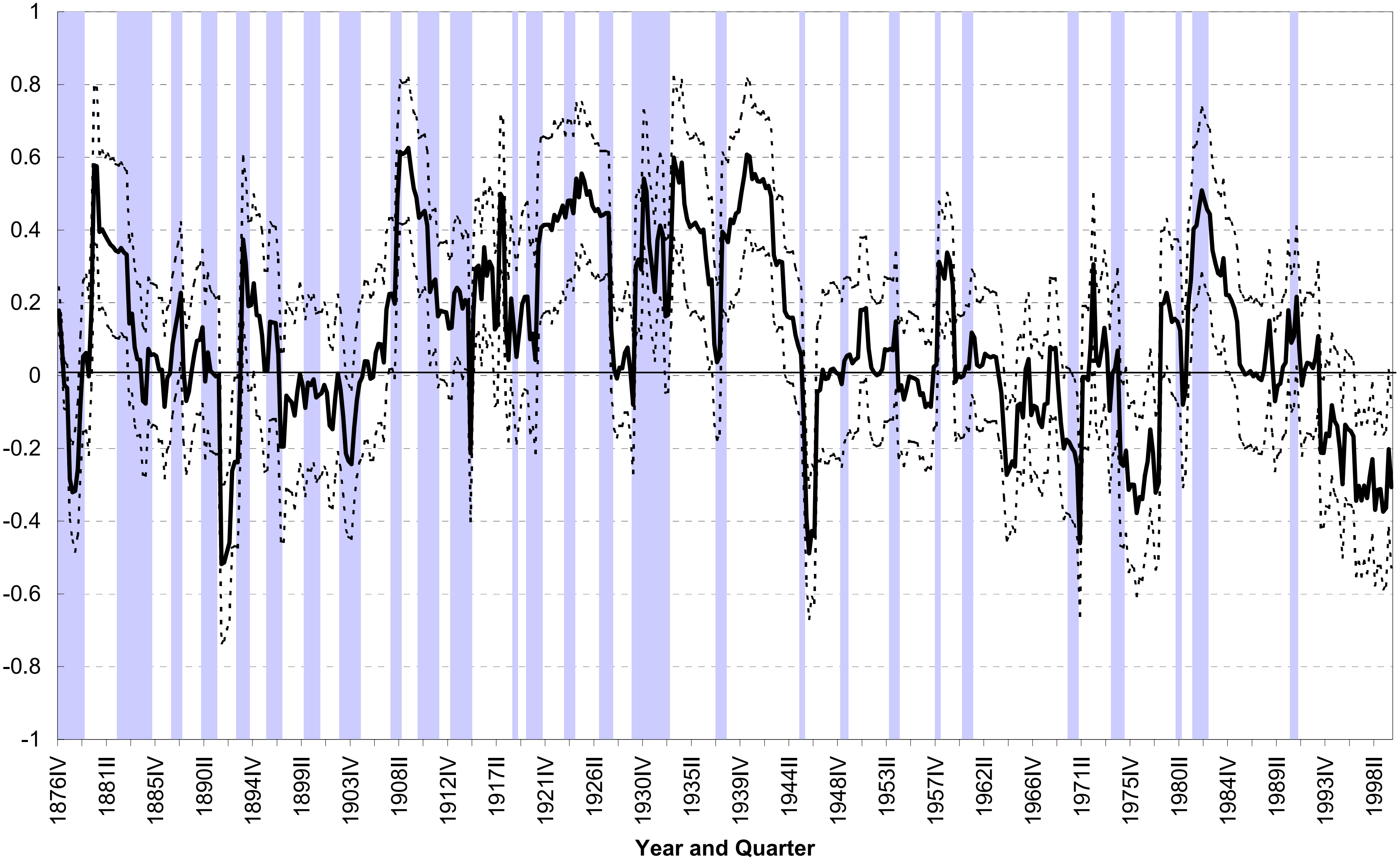

Figure 2
Values of Rho Significantly Different from Zero.
(Consecutive Signficant Values are Joined by a Line. Shaded areas are NBER periods of recession.)

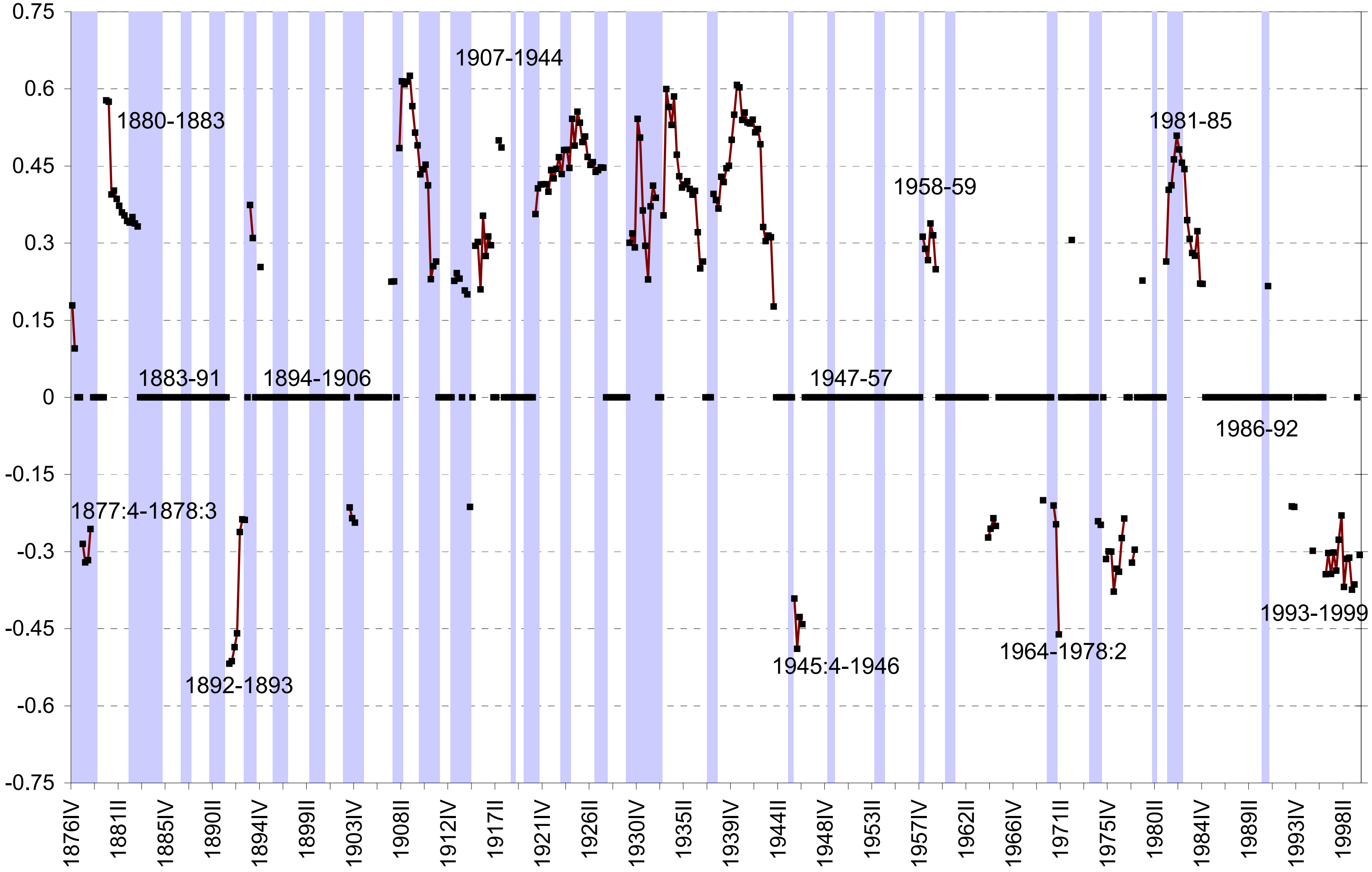

Figure 3
Price-Output Correlations at Different Forecast Horizons
(Shaded areas represent NBER periods of recession.)