Why Did the Sign of the Price-Output Correlation Change?

Evidence from a Structural VAR with GARCH Errors

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and

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(abstract)

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JEL Classifications:

E3 - Prices, Business Fluctuations, and Cycles
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It is generally agreed that the price-output correlation in the United States was positive prior to the Second World War, but became negative during the postwar period (at least by 1972). This paper offers evidence that the price-output correlation changed signs because of a decrease in the variability of aggregate demand. A structural VAR with bivariate GARCH (1,1) errors is used to estimate a times series of price-output correlations as well as of the conditional variances of the structural shocks to *AD* and *AS*. It is found that during the postwar period the price-output correlation is negative and significantly different from zero only when the standard deviation of the *AD* shock is less than that of the *AS* shock.

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1. Introduction

The literature on the cyclical behavior of the price level (or inflation) grew out of the belief that the correlation between fluctuations in the price level (or inflation) and output—the price-output correlation—has important implications for various theories of the business cycle. If changes in output within an aggregate demand-aggregate supply (*AD-AS*) framework are caused primarily by changes in *AD*, then one would expect the price-output correlation to be positive. On the other hand if changes in output are primarily caused by changes in *AS*, then one would expect the price-output correlation to be negative. One result of these observations is the existence of a rather extensive literature on the sign of the price-output correlation.

But empirical studies of the price-output correlation have raised more questions than they have answered. The primary reason for this and the focus of this paper is that the sign of the price-output correlation appears to be sample dependent. There is general agreement that the sign of the price-output correlation was positive prior to the Second World War, but became negative after 1972 (if not earlier). Among the studies that establish this are Kydland and Prescott (1990), Cooley and Ohanian (1991), Wolf (1991), Smith (1992), Backus and Kehoe (1992), Hartley (1999) and Cover and Hueng (2003).[1] This raises the important question, "Why has the sign of the price-output correlation changed?"

This paper offers an answer to this question by using a structural VAR to produce a time series of estimates of the variance-covariance matrix of the structural shocks to the *AD* and *AS*

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[1] This conclusion is robust to the method of detrending the data and also, as established by Smith (1992) and Backus and Kehoe (1992), holds across 10 industrialized nations. Cover and Hueng (2003) provide a summary of many previous studies on the sign of the price-output correlation, while Hartley (1999) shows that the signs and sizes of other correlation coefficients also are sample dependent.

curves. Within the model employed here two things can cause an initially positive price-output correlation to become negative: (1) A sufficient decline in the ratio of the standard deviation of the structural *AD* shock (σ_η) to the standard deviation of the structural *AS* shock (σ_ε); and (2) a sufficient decrease in the correlation coefficient between the two structural shocks (ρ). We find that the standard deviation ratio ($\sigma_\eta/\sigma_\varepsilon$) began a substantial decline during 1934, while the correlation coefficient (ρ) between the two structural shocks increased sharply in late 1925 and has remained relatively high ever since. Since the correlation coefficient has moved in a direction that would tend to increase the price-output correlation, it must be the decline in the standard deviation ratio that is responsible for the change in the sign of the price-output correlation. In particular, for the postwar period whenever the standard deviation ratio is less than one ($\sigma_\eta/\sigma_\varepsilon < 1$), the price-output correlation is negative.

Although determining the reason for the decline in the relative standard deviation of the structural shock to aggregate demand is beyond the scope of this paper, the most likely cause is an improved monetary policy. It is generally agreed that better macroeconomic policy (both monetary and fiscal) is at least partly responsible for the increased stability of the U. S. economy since the Second World War. Because it is possible for a poor monetary policy to increase the variability of *AD*, it follows that whenever the variability of *AD* declines monetary policy at least did not prevent the improvement. That is, even if the increased stability of *AD* comes out of changes in the private sector, monetary policy at least has not prevented the reduced variability of *AD* from being realized.

Our results are also important because they provide evidence that the existence of a negative price-output correlation is consistent with a model in which changes in *AD* can have a substantial effect on the economy. Our sample period is 1876:II-2004:IV and our structural

VAR assumes that the effect of any given structural *AD* or *AS* shock is the same throughout the sample. Hence we show that all that is required to the change of the sign of the price-output correlation is for the relative variabilities of the structural *AD* and *AS* shocks to change.

The remainder of this paper is outlined as follows. Section 2 uses a survey of the literature on the price-output correlation to set up the approach employed here. Section 3 presents and discusses the structural VAR that we employ. Section 4 presents and discusses estimates of the time-varying variance-covariance matrix (VCV) of the VAR residuals, while section 5 presents and discusses estimates of the VCV of the structural shocks to *AD* and *AS*. Section 6 closes the paper with some further discussion of the results and offers some conclusions.

2. The Price-Output Correlation.

The recent literature on the price-output correlation began with the presumption that the price-output correlation is positive, and that this fact itself rules out the empirical usefulness of real-business cycle models because such models imply a negative price-output correlation (Mankiw, 1989, p. 88).[2] Kydland and Prescott (1990) responded to this assertion by denying that the price-output correlation is positive; rather they presented evidence that the correlation is negative (at least since the Korean War) and suggested that this rules out demand-driven theories

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[2] Mankiw's (1989) argument was based on the widespread belief at that time that the price-output correlation is positive. Examples of writers who reported positive price-output correlations prior to 1990 include 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, interestingly enough, 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.

of the business cycle (Kydland and Prescott, 1990, p. 1).[3] This initial exchange of views on the importance of the price-output correlation turned out to be based on a naïve view of what determines its sign and size. The theoretical literature on the price-output correlation shows that it is possible to create a reasonable macroeconomic model in which the price-output correlation can be positive, negative or essentially zero.

This literature begins with models that show that Keynesian or demand-driven business-cycle models are not necessarily inconsistent with a negative price-output correlation. In particular, Chada and Prasad (1993), Ball and Mankiw (1994), Judd and Trehan (1995) and Rotemberg (1996) present models that yield a negative price-output correlation even though they do not contain a supply shock. Rather, in each case the existence of a demand shock in combination with price-level stickiness is able to produce a negative price-output correlation.[4]

The more recent literature includes papers that demonstrate that monetary policy could be the cause of the change in the sign of the price-output correlation. Pakko (2000) uses cross-spectral analysis to explore why the sign of the price-output correlation changed after 1972. The idea behind his paper is that it is possible for the contemporaneous correlation between two

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[3] As pointed out in footnote 2 above, Kydland and Prescott were not the first to argue that the price-output correlation is negative. Wolf (1991) reexamines their work and finds that the negative price-output correlation does not exist before 1973.

[4] Den Haan (2000) presents evidence that contradicts the applicability of these models to United States data. He uses a VAR to calculate the correlation between the forecast errors of price and output over several forecast horizons. He finds that at relatively short forecast horizons the correlation is either positive or zero, while at relatively long horizons it is negative. It takes a model with both demand and supply shocks to create this pattern. Spencer (1996) argues that the impulse response functions of a bivariate VAR (where the two variables are real GDP growth and inflation) are consistent with the general idea of these models.

variables to change as a result of changes in their co-spectrum at only a few frequencies. Pakko shows that this is indeed the case with the price-output correlation. The change in its sign after 1972 is the result of changes in its co-spectrum at relatively low frequencies, in particular those at the low end of the business-cycle range.[5] Pakko goes on to demonstrate that it takes a model with at least two shocks to produce a price-output co-spectrum similar to those obtained with U. S. data. To see if a change in monetary policy could possibly explain such a change in the price-output co-spectrum, Pakko simulates a monetary model with two shocks, a monetary shock (which acts like a demand shock) and an output shock. He finds that if monetary policy is implemented in a procyclical manner (such as might have been done during the prewar and interwar periods), it is possible for the price-output co-spectrum to be positive at all frequencies. On the other hand, if monetary policy is implemented by means of a constant growth rule, then the price-output co-spectrum is negative at all frequencies. Pakko's simulation results establish the possibility that a change in monetary policy is the reason why the U. S. price-output correlation changed signs.

A question that Pakko (2000) does not address is how an optimal monetary policy affects the price-output correlation. This is done by Flodén (2000), as well as Cover and Pecorino (2003). In Flodén's model the monetary authority minimizes a quadratic loss function that depends on deviations of output and the rate of inflation from their target values. Although for each of the parameter values he considers, the price-output correlation is negative, he does find that an increase in the monetary authority's preference for output stability increases the absolute

[5] The business cycle range of periods is from 6 to 32 quarters, so the low end of the range of business-cycle frequencies is 32 quarters.

value of this negative price-output correlation—that is, he finds that increasing the weight on output in a quadratic loss function reduces the price-output correlation[6].

Cover and Pecorino (2003) also consider the effect of policy on the price-output correlation when the monetary authority minimizes a quadratic loss function. In their model the monetary authority has the ability to observe shocks to *AD* and *AS* but only with observation errors. After observing the shocks (with error), the monetary authority chooses the optimal nominal rate of interest, which in their model has the effect of placing the *AD* curve in its optimal location. They find that the price-output correlation can be negative only if the monetary authority puts sufficient weight on both output and inflation in their loss function. That is, the closer the monetary authority is to being concerned with only one goal, the less likely it is that the price-output correlation will be negative.

Since in each of their models monetary policy affects the economy through its effect on *AD*, the papers by Flodén (2000), Pakko (2000) and Cover and Pecorino (2003) all suggest that the change in the sign of the price-output correlation could be the result of a change in the behavior of *AD*. Hence a reasonable place to start searching for an explanation of the change in the sign of the price-output correlation is to use an empirical model capable of estimating changes in the behavior of *AD*. We do this by combining the empirical model of Cover and Hueng (2003) with the structural VAR proposed by Cover, Enders and Hueng (2006).

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[6] Flodén considers the following loss function: $L = ½E[(\pi_t - \pi^*)^2 + \lambda(y_t - y^*)^2]$, where E is the expectation operator, π_t is inflation, y_t is output and π^* and y^* are the respective target values. He considers only two values of λ, 0.1 and 0.5. Although increasing λ from 0.1 to 0.5 is a small numerical increase, it represents decreasing the ratio of the weight on inflation to weight on output from 10 to 2.

Following den Haan (2000), Cover and Hueng (2003) define the price-output correlation to be the correlation coefficient between the unexpected changes in output and the price level obtained from estimating a VAR. Cover and Hueng, however, assume that the stochastic error term in their VAR follows a GARCH process. Using data for the period 1875:I-1999:IV, this allows them to obtain an estimate of a long time series of price-output correlations. They find that before 1945 the price-output correlation is usually positive or zero, while after the war it is usually negative or zero, results generally consistent with the other literature.

Cover, Enders and Hueng (2006) present a structural bivariate VAR (where the two variables are price and output) that allows one to obtain estimates of the VCV of the structural shocks to *AD* and *AS*. Unlike most other structural VAR's, their model does not require the structural shocks to be uncorrelated. In the next section we present their model and show that if it is estimated with a GARCH error term, it can be used to produce an estimate of the time series of the conditional VCV matrix of the structural shocks to *AD* and *AS*. We then show how changes in the VCV matrix of the structural *AD* and *AS* shocks affect the price-output correlation.[7]

3. The structural VAR

Consider the following model proposed by Cover, Enders, and Hueng (2006):

$$y_t^s = {}_{t\text{-}1}y_t + \alpha(P_t - {}_{t\text{-}1}P_t) + \varepsilon_t\,, \quad \alpha > 0; \tag{1}$$

$$(y_t + P_t)^d = {}_{t\text{-}1}(y_t + P_t)^d + \eta_t\,; \tag{2}$$

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[7] Lee (2004) implements a procedure similar to that of Cover and Hueng and then uses the Blanchard-Quah decomposition to obtain a time series of mutually uncorrelated structural *AD* and *AS* shocks. He then regresses his time-varying price-output correlation on these shocks and finds that *AD* shocks tend to make the correlation positive, while *AS* shocks tend to make it negative.

$$y_t^d = y_t^s; \tag{3}$$

where y_t and P_t respectively are output and the price level during period t, ${}_{t\text{-}1}y_t$ and ${}_{t\text{-}1}P_t$ are expected y_t and P_t given the information available at the end of period t-1, the superscripts s and d represent AS and AD, respectively; and ε_t and η_t respectively denote the serially uncorrelated structural AS and AD shocks. It is assumed that the variances of ε_t and η_t respectively are σ_ε^2 and σ_η^2, while their covariance is $\rho\sigma_\varepsilon\sigma_\eta$. ($\rho$ is the correlation coefficient between the two structural disturbances.) Equation (1) is a Lucas (1972) AS curve in which output increases in response to unexpected increases in the price level and positive realizations of the AS shock ε_t. Equation (2) is the aggregate demand relationship: nominal aggregate demand equals its expected value plus a random demand disturbance, η_t.

If equations (1) – (3) are solved for output and the price level the result is:

$$\begin{bmatrix} y_t \\ P_t \end{bmatrix} = \begin{bmatrix} {}_{t\text{-}1}y_t \\ {}_{t\text{-}1}P_t \end{bmatrix} + \begin{bmatrix} \dfrac{1}{1+\alpha} & \dfrac{\alpha}{1+\alpha} \\ \dfrac{-1}{1+\alpha} & \dfrac{1}{1+\alpha} \end{bmatrix} \cdot \begin{bmatrix} \varepsilon_t \\ \eta_t \end{bmatrix}. \tag{4}$$

If it is assumed that ${}_{t\text{-}1}y_t$ and ${}_{t\text{-}1}P_t$ are equal to linear combinations of their past observed values, then the system can be represented by a VAR, such as

$$\begin{bmatrix} y_t \\ p_t \end{bmatrix} = \begin{bmatrix} y_0 \\ p_0 \end{bmatrix} + \begin{bmatrix} a_{11}(L) & a_{12}(L) \\ a_{21}(L) & a_{22}(L) \end{bmatrix} \cdot \begin{bmatrix} y_t \\ p_t \end{bmatrix} + \begin{bmatrix} \dfrac{1}{1+\alpha} & \dfrac{\alpha}{1+\alpha} \\ \dfrac{-1}{1+\alpha} & \dfrac{1}{1+\alpha} \end{bmatrix} \begin{bmatrix} \varepsilon_t \\ \eta_t \end{bmatrix}; \tag{5}$$

which can be rewritten as:

$$\begin{bmatrix} y_t \\ p_t \end{bmatrix} = \begin{bmatrix} y_0 \\ p_0 \end{bmatrix} + \begin{bmatrix} a_{11}(L) & a_{12}(L) \\ a_{21}(L) & a_{22}(L) \end{bmatrix} \cdot \begin{bmatrix} y_t \\ p_t \end{bmatrix} + \begin{bmatrix} e_{yt} \\ e_{pt} \end{bmatrix}; \tag{6}$$

where $e_{y,t}$ and $e_{P,t}$ are the reduced-form VAR residuals and the $a_{ij}(L)$ are polynomials of order m in the lag operator, L, or $a_{ij}(L) = \sum_{k=1}^{m} a_{ij} L^k$. The implied relationship between the VAR residuals and the structural innovations in this *AD*-*AS* model is:

$$\begin{bmatrix} e_{yt} \\ e_{Pt} \end{bmatrix} = \begin{bmatrix} \frac{1}{1+\alpha} & \frac{\alpha}{1+\alpha} \\ \frac{-1}{1+\alpha} & \frac{1}{1+\alpha} \end{bmatrix} \cdot \begin{bmatrix} \varepsilon_t \\ \eta_t \end{bmatrix}. \tag{7}$$

Equation (7) can be used to derive the following relationship between the VCV matrix of the VAR residuals and the VCV matrix of the structural *AD* and *AS* shocks:

$$\begin{bmatrix} \mathrm{var}(e_y) & \mathrm{cov}(e_y, e_p) \\ \mathrm{cov}(e_y, e_p) & \mathrm{var}(e_p) \end{bmatrix} = \begin{bmatrix} \frac{1}{1+\alpha} & \frac{\alpha}{1+\alpha} \\ \frac{-1}{1+\alpha} & \frac{1}{1+\alpha} \end{bmatrix} \cdot \begin{bmatrix} \sigma_\varepsilon^2 & \rho\sigma_\varepsilon\sigma_\eta \\ \rho\sigma_\varepsilon\sigma_\eta & \sigma_\eta^2 \end{bmatrix} \cdot \begin{bmatrix} \frac{1}{1+\alpha} & \frac{-1}{1+\alpha} \\ \frac{\alpha}{1+\alpha} & \frac{1}{1+\alpha} \end{bmatrix}. \tag{8}$$

The left-hand side of equation (8) is the VCV matrix of the VAR residuals and the middle term on the right-hand side is the VCV matrix of the structural *AD* and *AS* shocks. The essential feature of the above model that makes it different from previous studies that employ a structural VAR is that the VCV matrix of the structural *AD* and *AS* shocks on the right-hand-side of (8) is not assumed to be an identity matrix (or even a diagonal matrix), rather the structural *AD* and *AS* shocks are allowed to be contemporaneously correlated.

Equation (8) represents a system of three equations (two variance equations and a covariance equation) with four unknowns (α, σ_ε, σ_η, and ρ). Cover, Enders, and Hueng show that the Blanchard and Quah (1989) restriction—the assumption that the structural aggregate demand shock (η_t) has no long-run effect on output—can be used to identify α, the slope of the short run *AS* curve.[8] They therefore are able to use the three equations in (8) to estimate the VCV of the structural *AD* and *AS* shocks.

This paper modifies the procedure of Cover, Enders and Hueng by assuming that the error terms in equation (6) follow a GARCH (1, 1) process. This allows estimation of a time series of conditional values for the VAR VCV (the left hand side of equation (8)), which in turn allows for the straightforward calculation of a time series of conditional values for the VCV of the structural *AD* and *AS* shocks.

To see how this is done, let $\boldsymbol{X_t} = [\, y_t \;\; P_t \,]'$, $\boldsymbol{e_t} = [e_{y,t} \;\; e_{P,t}]'$, $\boldsymbol{B_0} = [y_0 \;\; P_0]'$,

$\boldsymbol{B_j} = \begin{bmatrix} a_{11}(j) & a_{12}(j) \\ a_{21}(j) & a_{22}(j) \end{bmatrix}$ and rewrite equation (6) in the following way:

$$\boldsymbol{X_t} = \boldsymbol{B_0} + \sum_{j=1}^{m} \boldsymbol{B_j X_{t-j}} + \boldsymbol{e_t}\,; \tag{9}$$

$$\boldsymbol{H_t} = \boldsymbol{\Gamma}' \cdot \boldsymbol{\Gamma} + \boldsymbol{G}' \cdot \boldsymbol{H_{t-1}} \cdot \boldsymbol{G} + \boldsymbol{A}' \cdot \boldsymbol{e_{t-1}} \cdot \boldsymbol{e}'_{t-1} \cdot \boldsymbol{A}\,; \tag{10}$$

where $\boldsymbol{e}_t \mid \boldsymbol{\Omega}_{t-1} \sim N(0, \boldsymbol{H_t})$, $\boldsymbol{\Omega}_{t-1}$ is the information set at time $t-1$, $\boldsymbol{\Gamma}$ is upper triangular, $\boldsymbol{G}$ and $\boldsymbol{A}$ are diagonal, and the product of the diagonal elements of $\boldsymbol{G}$ is restricted to be positive. Equations (9) and (10) form the bivariate VAR(m)-GARCH(1,1) model employed by Cover and Hueng (2003) to estimate a time series of conditional price-output correlations for the United States.

In (10) $\boldsymbol{H_t}$ is the conditional VCV of the VAR error terms for period t. Because α is not time-varying in this model, if the left-hand side of (8) is replaced with $\boldsymbol{H_t}$, the result will be the

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[8] It is worth noting that the value of α depends only on the estimated autoregressive coefficients in the VAR (the $a_{ij}(L)$), so its value is not time-varying.

following equation that shows the relationship between the conditional VCV matrix of the VAR error terms for period *t* and the conditional VCV matrix of the structural shocks for period *t*:

$$H_t = \begin{bmatrix} \mathrm{var}_{t-1}\left(e_{y,t}\right) & \mathrm{cov}_{t-1}\left(e_{y,t}, e_{P,t}\right) \\ \mathrm{cov}_{t-1}\left(e_{y,t}, e_{P,t}\right) & \mathrm{var}_{t-1}\left(e_{y,t}\right) \end{bmatrix} =$$

$$\begin{bmatrix} \frac{1}{1+\alpha} & \frac{\alpha}{1+\alpha} \\ \frac{-1}{1+\alpha} & \frac{1}{1+\alpha} \end{bmatrix} \cdot \begin{bmatrix} \sigma^2_{\varepsilon,t|t-1} & \rho_{t|t-1}\sigma_{\varepsilon,t|t-1}\sigma_{\eta,t|t-1} \\ \rho_{t|t-1}\sigma_{\varepsilon,t|t-1}\sigma_{\eta,t|t-1} & \sigma^2_{\eta,t|t-1} \end{bmatrix} \cdot \begin{bmatrix} \frac{1}{1+\alpha} & \frac{-1}{1+\alpha} \\ \frac{\alpha}{1+\alpha} & \frac{1}{1+\alpha} \end{bmatrix}. \qquad (11)$$

4. Estimated Conditional VCV Matrix of VAR Errors.

The U.S. quarterly time series data on real GDP and nominal GDP are taken from the United States Department of Commerce for the period 1947:I-2004:IV and from Balke and Gordon (1986) for the period 1875:I-1947:I. The price level for each quarter is defined as the implicit price deflator implied by the values of real and nominal GDP. The data sets are spliced together so that the quarterly growth rates of real GDP and the implicit price deflator through the first quarter of 1947 are equal to those in the Balke-Gordon data, while those beginning with the second quarter of 1947 are equal to those in the Department of Commerce data. Standard Augmented Dickey-Fuller tests of the logarithms of real GDP and the GDP deflator indicated that both variables are difference stationary. Hence, the variables employed in the VAR are the first-difference of log real GDP and the first-difference of the log GDP deflator.

Before estimation of (9) and (10) the log likelihood ratio test used in Sims (1980) was employed to determine that the optimal lag length *m* in (9) is four lags. After the joint estimation of equations (9) and (10) by maximum likelihood a time series of estimated correlation coefficients was calculated from the time series of $\hat{\boldsymbol{H}}_t$, the estimated conditional VCV matrix of the VAR error terms for period *t*. For inference, a Monte Carlo simulation was used to obtain confidence intervals (see Cover and Hueng (2003) for details). The estimated correlation coefficients are plotted in Figure 1, where the shaded areas indicate statistically significant estimates, i.e., the estimates that fall outside the $\pm$ one-standard deviation bands.

There are two important things to note about Figure 1. The first and perhaps most obvious one is that significantly negative price-output correlations are rare before 1945, while significantly positive price-output correlations are rare after 1945. Somewhat less obvious and not emphasized in the literature prior to Cover and Hueng (2003) is that Figure 1 shows that the price-output correlation is not significantly different from zero most of the time both before and after 1944. This is summarized in Table 1 which shows that for over one-half of the quarters—both pre and postwar—the price-output correlation is not significantly different from zero.

Table 1

Share of quarters with a significantly positive and a significantly negative price-output correlation.

Period	Share of quarters positive and significant	Share of quarters negative and significant	Share of quarters not significantly different from zero at 5% level
1876:II -2004:IV	25.6%	21.6%	52.8%
1876:II - 1944:IV	44.4%	4.0%	51.6%
1945:I – 2004:IV	4.2%	41.7%	54.2%
1919:I – 1939:IV	70.2%	0.0%	29.8%
1984:I - 2004:IV	0.0%	59.5%	40.5%

As is also shown in Table 1 there is one long sub-period, 1919-1939 in which positive correlations are more common than essentially zero correlations, as well as one long period, 1984-2004, in which negative correlations are more common than essentially zero correlations. Hence, consistent with the findings in Cover and Hueng, it must be concluded that that the estimated correlation is usually positive or essentially zero before 1945 and usually negative or essentially zero after 1945.

Before beginning the discussion of the estimation of the structural model, it is useful to examine the behavior of the time series of $var_{t-1}(e_{y,t})$ and $var_{t-1}(e_{P,t})$, the conditional variances of the unexpected changes in output growth and inflation. The natural logarithms of their standard deviations are presented in Figure 2. What is striking about Figure 2 is how the standard deviations of unexpected output growth and inflation have changed over time. In particular, before the mid-1920's the conditional standard deviation of unexpected output growth and unexpected inflation were approximately the same (though the latter is more variable), while after 1924 that of unexpected inflation is always lower than that of unexpected output growth (although it remains more variable). Both conditional standard deviations exhibit long secular declines, but the decline for the conditional standard deviation of unexpected inflation begins about one decade earlier and is deeper than the decline in the conditional standard deviation of output growth.[9]

5. Estimated Conditional VCV Matrix of Structural Shocks.

Once $\hat{\boldsymbol{H}}_t$ is obtained from estimating (9) and (10), equation (11) is a matrix equation that includes three equations but four unknowns. But as discussed above, imposing the Blanchard-Quah restriction yields an estimate of α, the slope of the short-run *AS* curve. For our sample and specification the estimated value of α is $\hat{\alpha} = 0.706$. Given this value of $\hat{\alpha}$, equation (11) was

[9] The decline in the standard deviation of output growth that Figure 2 suggests begins in 1983 has been noted in the literature (e.g., McConnell and Perez-Quiros, 2000), but the decline beginning in the mid-1930's shown in Figure 2 has not (though there is general agreement that output during the postwar period is more stable than that during the interwar period). Finally, the decline in the variability of inflation has not been emphasized.

used to obtain a time series of the conditional VCV matrix of the structural shocks. Before presenting the estimates of the structural VCV it is useful to consider the following equation for the covariance between unexpected output and unexpected inflation obtained from (11):

$$cov(e_y, e_p) = \frac{\sigma_\varepsilon^2}{(1+\alpha)^2}\left[\alpha\left(\frac{\sigma_\eta}{\sigma_\varepsilon}\right)^2 + \rho(1-\alpha)\left(\frac{\sigma_\eta}{\sigma_\varepsilon}\right) - 1\right]. \qquad (12)$$

Equation (12) shows how the VCV matrix of the structural shocks determines the sign of the price-output correlation. Given that $\hat{\alpha} > 0$, if $\hat{\rho} > 0$ (which is the case for 83% of our observations), the first two terms on the RHS of (12) are positive and therefore imply that the higher is the standard deviation of the structural *AD* shock relative to the standard deviation of the structural *AS* shock, the more likely it is that the price-output correlation is positive. It is also the case that the larger is $\hat{\rho}$ (given that $\hat{\alpha} < 1$), then the more likely it is that the price-output correlation is positive. Therefore equation (12) suggests that an examination of the time series behavior of the structural standard deviation ratio, ($\sigma_\eta/\sigma_\varepsilon$), and the structural correlation coefficient ρ will yield insight into why the price-output correlation has changed.

Finally, equation (12) implies that for any given values of α and ρ there is always a standard deviation ratio, ($\sigma_\eta/\sigma_\varepsilon$), which causes the price-output correlation to equal zero. The preponderance of quarters in our sample with a conditional price-output correlation not significantly different from zero (as shown above in Table 1) suggests that the standard deviation ratio has often been very close to this value.[10]

[10] The value of the standard deviation ratio that causes the price-output correlation to equal zero for $\alpha = 0.706$ and various values of ρ is shown in the following table:

ρ	0.6	.75	0.9
$\sigma_\eta/\sigma_\varepsilon$	1.154	1.145	1.135

Figure 3 presents the estimated conditional standard deviations of the *AD* shock ($\hat{\sigma}_{\eta,t|t-1}$) and the *AS* shock ($\hat{\sigma}_{\varepsilon,t|t-1}$), Figure 4 presents their ratio ($\frac{\hat{\sigma}_{\eta,t|t-1}}{\hat{\sigma}_{\varepsilon,t|t-1}}$), and Figure 5 presents the estimates of their conditional correlation coefficients ($\hat{\rho}_{t|t-1}$). Note that the shaded areas, as in Figure 1, are the periods in which the estimated price-output correlations are statistically significant.

Figure 3 shows that the conditional standard deviations of the structural *AD* and *AS* shocks tend to move in the same direction. They both begin a secular decline around 1935. Although the decline continues throughout the remainder of the sample, during the period from 1971 through 1982 they both increase slightly. Prior to 1945 the conditional standard deviation of the structural *AD* shock is usually larger than that of the structural *AS* shock. That is, when the price-output correlation is positive and significantly different from zero, the structural *AD* shock is larger than the structural *AS* shock. This is reaffirmed by noting that during the only postwar quarters (1950:IV, 1972:II, and 1981:II-1983:I) with a positive and significantly different from zero price-output correlation, the conditional variance of the structural *AD* shock is greater than that of the structural *AS* shock. After 1945 the two conditional standard deviations are approximately the same size. However, during periods when the price-output correlation is negative and significantly different from zero, the conditional standard deviation of the structural *AS* shock is greater than that of the structural *AD* shock.

Figure 4 presents the ratio of the conditional standard deviations, ($\sigma_\eta/\sigma_\varepsilon$). It shows clearly that the periods of negative and significant price-output correlations during the postwar period are also periods when the standard deviation ratio is less than one. This decline in the standard deviation ratio must be the cause of the change in the sign of the price-output correlation because, as shown in Figure 5, the conditional correlation coefficient between the two structural disturbances is generally higher during the postwar period. During the period 1876:II-1922:IV the correlation coefficient between the structural *AD* and *AS* disturbances is highly variable and just as likely to be negative as positive.[11] But after 1922 the estimated correlation is positive every quarter and clearly less variable after 1930. After 1930 the estimated conditional correlation is above 0.8 for 65% of the observations, and above 0.6 for 94% of the observations.

To conclude this section, as is shown by equation (12), an increase in the conditional correlation coefficient, $\hat{\rho}$, makes it more likely that the price-output correlation is positive. Since the high values of $\hat{\rho}$ during the postwar period would tend to cause the price-output correlation to be positive, it must be concluded that the change in the structural standard deviation ratio is the primary cause of the negative price-output correlation during the postwar period.

6. Summary and Conclusions

This paper examines why the sign of the price-output correlation has changed from being generally positive prior to the Second World War to generally negative during the postwar period. Since some prominent writers have argued that the sign of the price-output correlation can rule out models that predict the wrong sign, an explanation of the change in the sign of the price-output correlation is important.

Using an empirical procedure similar to that of Cover and Hueng (2003) this paper shows that the price-output correlation usually has been insignificantly different from zero, but when significant the sign was usually positive before the war and negative during the postwar period.

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[11] During the period 1876II-1922IV the estimated conditional correlation coefficient between the structural

By employing a GARCH version of the structural VAR model of Cover, Enders and Hueng (2006) we estimated a time series of the standard deviations of the structural shocks to *AD* and *AS*. We found that a decline in the ratio of the *AD* standard deviation to the *AS* standard deviation is the primary reason for the change in the sign of the price-output correlation.

Our results imply that *AD* has become a less important source of fluctuations in output during the postwar period even if there has been no change in the structure of the economy. That is, *AD* appears to be less important only because it has been relatively less variable than *AS*. Since it is always possible for a poor monetary policy to make *AD* highly variable, it must be concluded that even if the increased stability in *AD* is not the direct result of a better monetary policy, monetary policy at least has not been harmful. That is, our results are consistent with the idea that an improved monetary policy is the primary cause of the change in the sign of the price-output correlation as well as the idea that improved macroeconomic policy in general has helped to stabilize the economy during the postwar era.

Our results also suggest that both policymakers and economists should exercise caution when using estimated correlations in available data to evaluate alternate macroeconomic models. Once the effect of monetary policy is taken into account, the sign and size of the price-output correlation places no restrictions on the general form of a suitable macroeconomic model.

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shocks was negative for 47% of the observations.

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Figure 1
Estimated Price-Output Correlations: 1876II - 2004IV

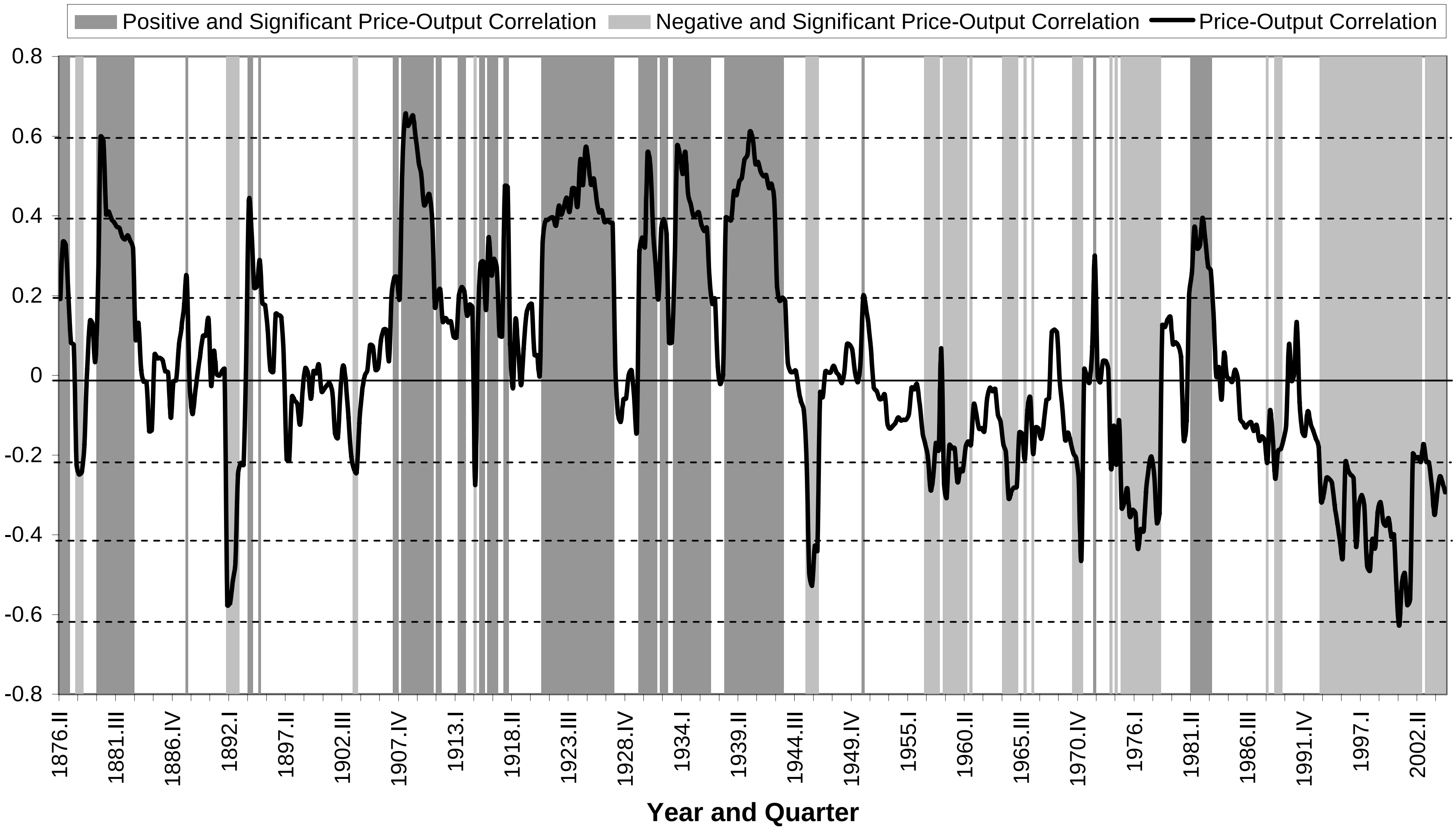

Figure 2
Standard Deviations of Unexpected Changes in Output Growth and Inflation

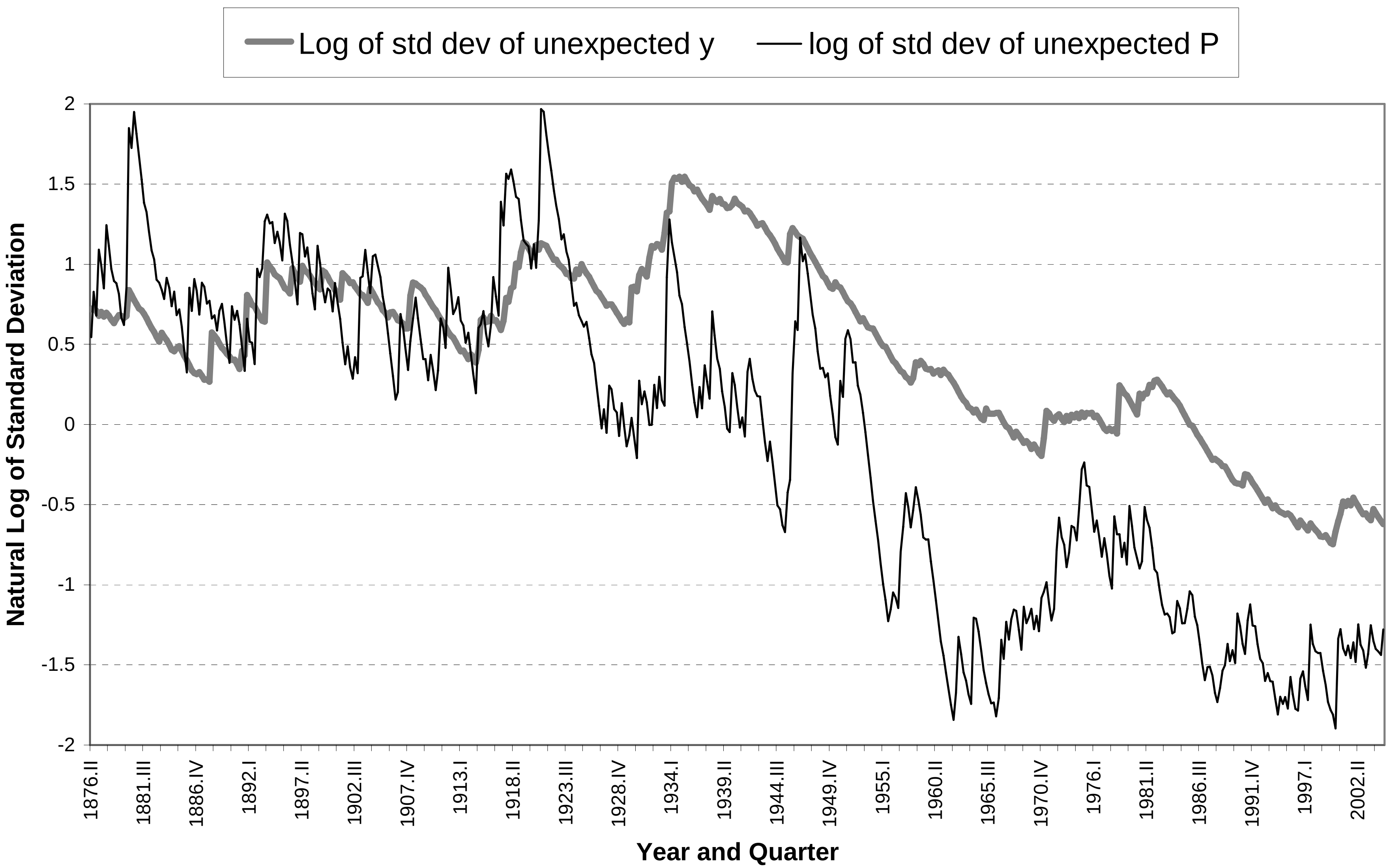

Figure 3
Standard Deviations of Structural AD and AS Shocks

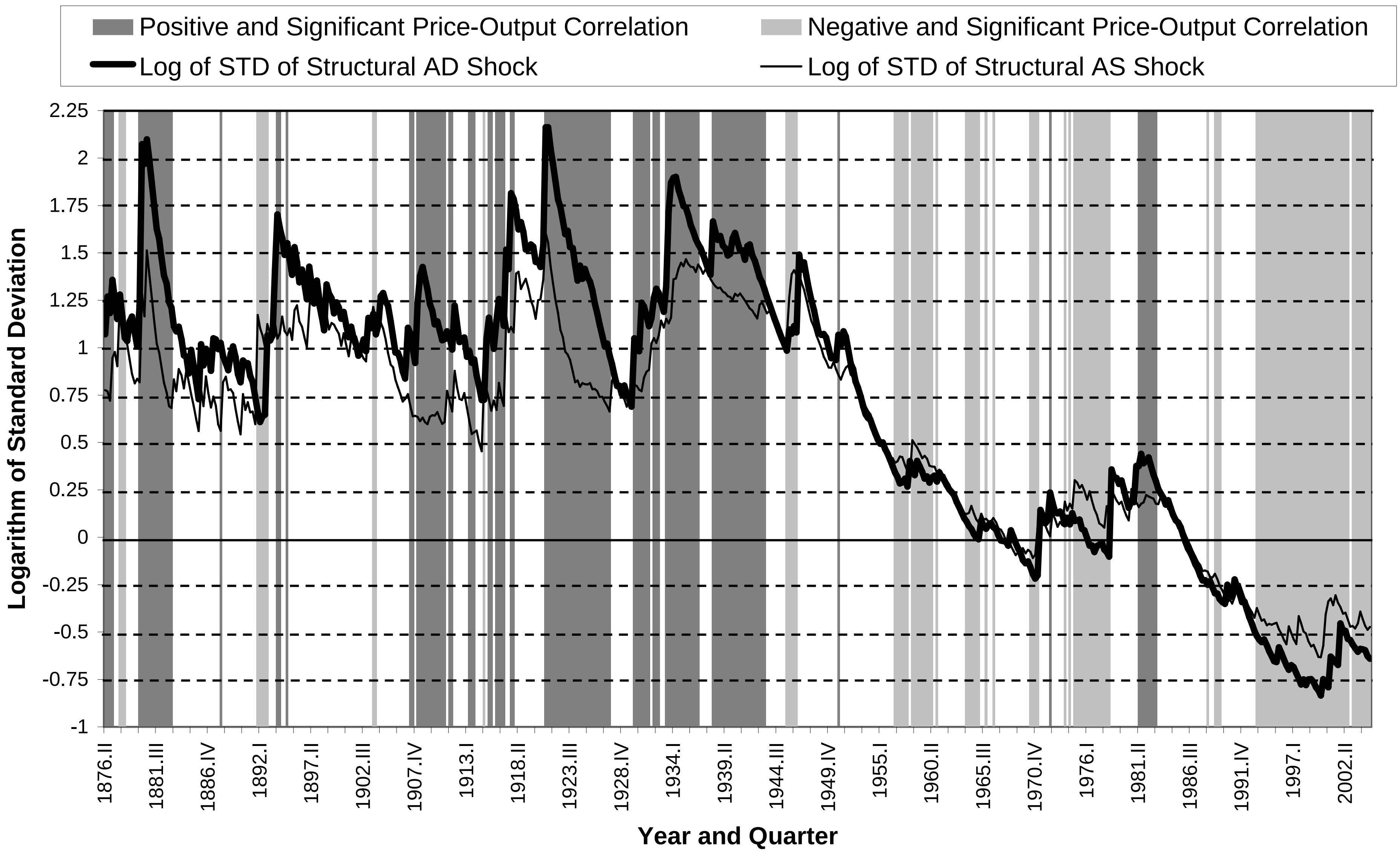

Figure 4
Ratio of Standard Deviation of Structural AD Shock to Structural AS Shock

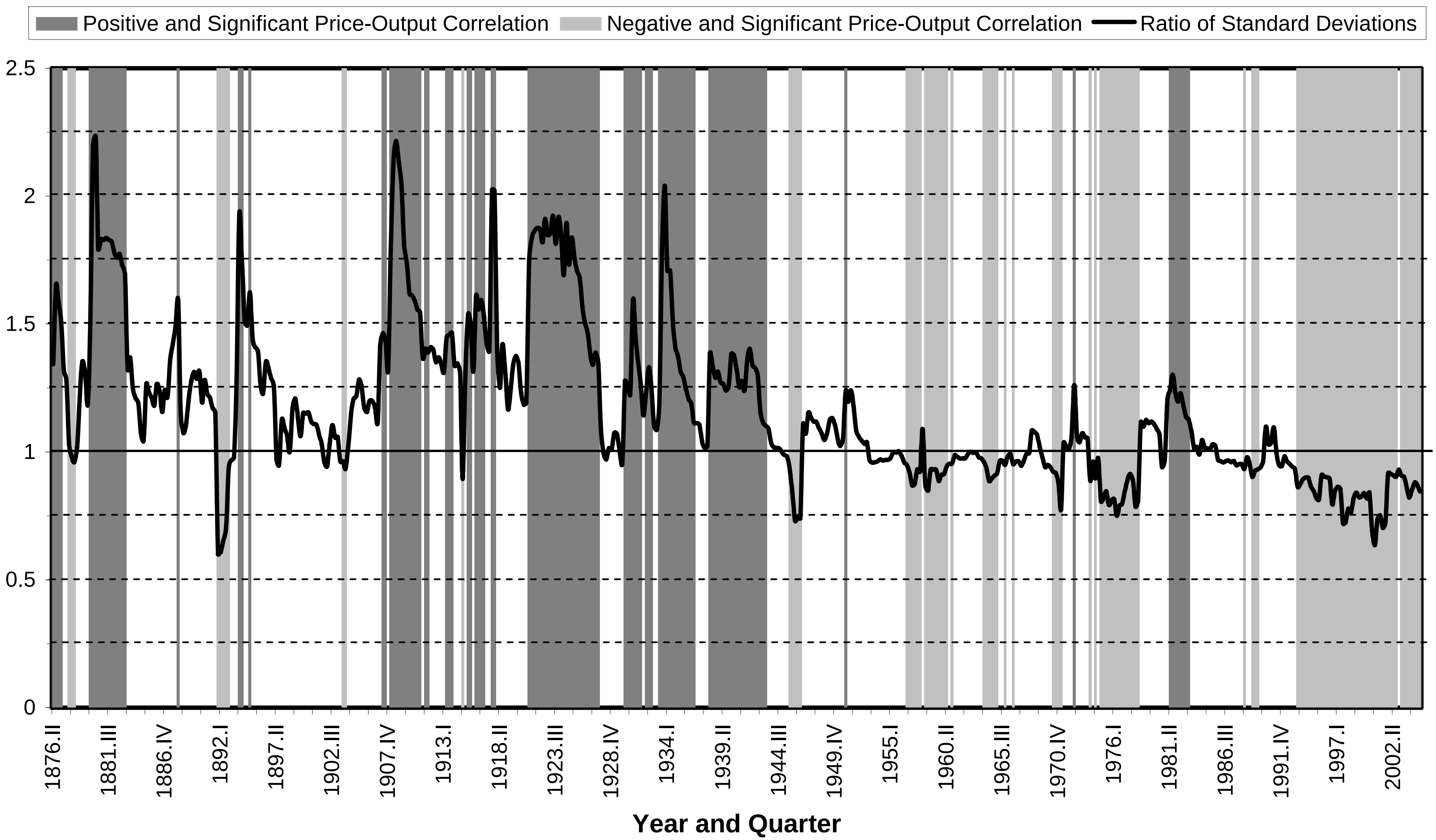

Figure 5
Conditional Correlation Coefficient between Structural AD and AS Shocks