Unit Labor Costs and the Price Level

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A popular theoretical model of the inflation process is the expectations-augmented Phillips-curve model. According to this model, prices are set as markup over productivity-adjusted labor costs, the latter being determined by expected inflation and the degree of demand pressure.[1] It is assumed further that expected inflation depends upon past inflation. This model thus implies that productivity-adjusted wages and prices are causally related with feedbacks running in both directions.

In this article, I investigate empirically the causal relationship between prices and productivity-adjusted wages (measured by unit labor costs) using cointegration and Granger-causation techniques.[2] In my recent paper, Mehra (1991), I used similar techniques[3] to show that inflation and growth in unit

[1] This version has been closely associated with the work of Gordon (1982, 1985, 1988) and differs from the original Phillips-curve model. The latter was formulated as a wage equation relating wage inflation to the unemployment gap.

[2] Let X_{1t}, X_{2t}, and X_{3t} be three time series. Assume that the levels of these time series are nonstationary but first differences are not. Then these series are said to be cointegrated if there exists a vector of constants $(\alpha_1, \alpha_2, \alpha_3)$ such that $Z_t = \alpha_1 X_{1t} + \alpha_2 X_{2t} + \alpha_3 X_{3t}$ is stationary. The intuition behind this definition is that even if each time series is nonstationary, there might exist linear combinations of such time series that are stationary. In that case, multiple time series are said to be cointegrated and share some common stochastic trends. Moreover, if series are cointegrated, then some series must adjust in the short run so as to maintain equilibrium among multiple series. That implies the presence of short-run feedbacks (and hence Granger-causality) among these series.

[3] The statistical inference in most of the empirical work prior to Mehra (1991) has often been conducted under the assumption that wage and price series contain deterministic trends. Recent evidence has called this assumption into question and has shown that the trend components of several of these time series also contain stochastic components (Nelson and Plosser 1982). A misspecification of trend components can lead to incorrect tests of hypotheses. Mehra (1991) therefore employed recent techniques to investigate trends in wage and price series and used the analysis to determine the nature of causal structure between prices and unit labor costs.

labor costs are correlated in the long run and that the presence of this correlation appears to be due to Granger-causality running from inflation to growth in unit labor costs, not the other way around. The results presented there indicate that the "price markup" hypothesis is inconsistent with the data and that growth in unit labor costs does not help predict the future inflation rate.

This article examines the robustness of the conclusions in Mehra (1991) to changes in the measure of the price level, the sample period, and unit root-cointegration test procedures used there. In particular, the price series used in Mehra (1991) is the fixed-weight GNP deflator that covered the sample period 1959Q1 to 1989Q3; the test for cointegration used is the two-step procedure given originally in Engle and Granger (1987); and the stationarity of data is examined using Dickey-Fuller unit root tests. This article considers an additional price measure, the consumer price index, which covers consumption goods and services bought by urban consumers. In contrast, the implicit GNP deflator, the other price measure used here, covers prices of consumption, investment, government services, and net exports. Since the consumer price index is also a widely watched measure of inflation pressures in the economy, the article examines whether the causal relationships found between the general price level and unit labor costs carry over to consumer prices.

In my earlier empirical work (1991), I used Dickey-Fuller unit root tests to determine whether the relevant series contain stochastic or deterministic trends. Recently, some authors including Dejong et al. (1992) have shown that Dickey-Fuller tests have low power in distinguishing between these two alternatives. These studies suggest that economists should supplement unit root tests by tests of trend stationarity. Thus, a series now is considered having a unit root if two conditions are met: (1) the series has a unit root by Dickey-Fuller tests and (2) it is not trend stationary by tests of trend stationarity. Furthermore, the test for cointegration recently proposed by Johansen and Juselius (1990) overcomes several pitfalls associated with the Engle-Granger test for cointegration.[4] This article employs these additional, refined cointegration-stationarity tests to determine the stationarity of data and to study the nature of the causal structure between the general price level and unit labor costs.

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[4] The Engle-Granger test for cointegration is implemented by regressing one series on the other remaining series and then testing whether the residuals from that regression are stationary or not. If the residuals are stationary, then the multiple time series are said to be cointegrated. This test has several shortcomings: (1) the test results are sensitive to the particular series chosen as the dependent variable; (2) the test cannot tell whether the number of cointegrating relationships is one or more than one; and (3) tests of hypotheses in the cointegrating vectors cannot be carried out because estimated coefficients have unknown nonstandard distributions. In contrast, the test proposed in Johansen and Juselius (1990) does not have any of the aforementioned problems. Their test procedure enables one to test directly for the number of cointegrating vectors and provides at the same time the maximum likelihood estimates of the cointegrating vectors. Tests of hypotheses in such estimated cointegrating vectors can be easily carried out. Lastly, the test results are not sensitive to the particular normalization chosen.

The empirical evidence reported here indicates that wage and price series contain stochastic, not deterministic, trends and that long-run movements in prices are correlated with long-run movements in unit labor costs. That is, the wage and price series used here are cointegrated as discussed in Engle and Granger (1987). This result holds whether the particular price series used is the implicit GDP deflator or the consumer price index.

Tests of Granger-causality presented here indicate that short-run movements in prices and unit labor costs are also correlated, with Granger-causality running one way from prices to unit labor costs when the price series used is the implicit GDP deflator. Test results with the consumer price index, however, are consistent with the presence of bidirectional feedbacks between prices and unit labor costs.

The empirical work here supports and extends the results in Mehra (1991). Though the cointegration test procedures, the sample periods, and the general price-level series used in these studies differ, both studies indicate that the "price markup" hypothesis is inconsistent with the data when the price series used measures the general price level. The additional results here, however, indicate such is not the case when the price series used is less broadly measured by the consumer price index. Thus, movements in unit labor costs help predict movements in consumer prices, but not in the general price level.

The plan of this article is as follows. Section 1 presents a Phillips-curve model of the inflation process and discusses its implications for the relationship between wages and prices. It also discusses how tests for cointegration and Granger-causality can be used to examine such wage-price dynamics. Section 2 presents the empirical results, and Section 3 contains concluding observations.

1. THE MODEL AND THE METHOD

The Phillips-Curve Model

The view that systematic movements in wages and prices are related derives from the expectations-augmented Phillips-curve model of the inflation process. Consider the price and wage equations that typically underlie such Phillips-curve models described in Gordon (1982, 1985) and Stockton and Glassman (1987):

$$\Delta P_t = h_0 + h_1 \Delta(w - q)_t + h_2 \chi_t + h_3 Sp_t \tag{1}$$

$$\Delta(w - q)_t = k_0 + k_1 \Delta P_t^e + k_2 \chi_t + k_3 Sw_t \tag{2}$$

$$\Delta P_t^e = \sum_{j=1}^{n} \lambda_j \Delta P_{t-j}, \tag{3}$$

where all variables are in natural logarithms and where P_t is the price level, w_t is the wage rate, q_t is labor productivity, χ_t is a demand-pressure variable, P_t^e is the expected price level, Sp_t represents supply shocks affecting the price equation, Sw_t represents supply shocks affecting the wage equation, and Δ is the first-difference operator. Equation (1) describes the price markup behavior. Prices are marked up over productivity-adjusted labor costs $(w - q)$ and are influenced by cyclical demand (χ) and the exogenous relative price shocks (Sp). This equation implies that productivity-adjusted wages determine the price level, given demand pressure. Equation (2) is the wage equation. Wages are assumed to be a function of cyclical demand (χ) and expected price level, the latter modeled as a lag on past prices as in equation (3). The wage equation, together with equation (3), implies that wages depend upon past prices, ceteris paribus.[5]

The price and wage behavior described above suggests that long-run movements in wages and prices must be related. In fact, some formulations of (1) and (2) predict that these two variables would grow at similar rates in the long run.[6] Furthermore, if one allows for short-run dynamics in such behavior, the analysis presented above would also suggest that past changes in wages and prices should contain useful information for predicting future changes in those same variables, ceteris paribus. These implications can be examined easily using tests for cointegration and Granger-causality between wage and price series.

Tests for Cointegration and Granger-Causality

If wage and price series have stochastic trends that move together, then the two time series should be cointegrated as discussed in Granger (1986). Thus, the long-run comovement of wages and prices is examined using the test for cointegration given in Johansen and Juselius (1990). The test procedure,

[5] The price and wage equations used here should be viewed as the reduced form equations. Price behavior as characterized in equation (1) is based on a markup model of pricing by firms. Nordhaus (1972) shows such pricing could be derived from optimizing behavior in which the technology is characterized by a Cobb-Douglas production function. Gordon (1985), on the other hand, derives a wage equation like (2) from an explicit model of labor demand and supply in which the wage rate adjusts in response to any change in the size of the gap between the two.

[6] For example, as indicated in footnote 5, the markup model of pricing behavior characterized in equation (1) is consistent with optimizing behavior in which the technology is characterized by a Cobb-Douglas production function. Given the additional assumptions of constant returns and the constant relative price of capital, the production environment implies a long-term coefficient of unity on unit labor costs in the price equation (1), $h_1 = 1$. That result indicates that prices and wages would grow at the same rate in the long run. Alternatively, the natural rate hypothesis, if valid in the long run, would indicate that the sum of the coefficients on past prices in (2) should be one, $k_1 \sum_{j=1}^{n} \lambda_j = 1$. That result also would indicate wages and prices grow at similar rates in the long run.

denoted hereafter as the JJ procedure, consists of estimating a VAR model that includes differences as well as levels of nonstationary time series. The matrix of coefficients that appear on levels of these time series contains information about the long-run properties of the model.

To explain the model, let X_t be a vector of time series on prices and wages. The VAR model is

$$X_t = \Pi_1 X_{t-1} + \Pi_2 X_{t-2} + \Pi_k X_{t-k} + \epsilon_t, \tag{4}$$

where $\Pi_i, i = 1, \ldots k$, is a matrix of coefficients that appear on X_{t-i}. Under the hypothesis that the series in X_t are difference stationary, it is convenient to transform (4) in a way that it contains both levels and first differences of the time series in X_t. That transformation is shown in (5).

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1} \Delta X_{t-k-1} + \Pi X_{t-k} + \epsilon_t, \tag{5}$$

where $\Gamma_i, i = 1, \ldots k-1$, and Π are matrices of coefficients that appear on first differences and levels of the time series in X_t. The component ΠX_{t-k} in (5) gives different linear combinations of levels of the time series in X_t. Thus, the matrix Π contains information about the long-run properties of the model. When the matrix's rank[7] is zero, equation (5) reduces to a VAR in first differences. In that case, no series in X_t can be expressed as a linear combination of other remaining series. That result indicates that there does not exist any long-run relationship between the series in the VAR. On the other hand, if the rank of Π is one, then there exists only one linearly independent combination of series in X_t. That result indicates that there exists a unique, long-run (cointegrating) relationship between the series. When the rank is greater than one, then there is more than one cointegrating relationship among the elements of X_t.

Two test statistics can be used to evaluate the number of the cointegrating relationships. The trace test examines the rank of Π matrix and the hypothesis that rank $(\Pi) \leq r$ is tested, where r represents the number of cointegrating vectors. The maximum eigenvalue test tests the null that the number of cointegrating vectors is r given the alternative of $r + 1$ vectors. The critical values of these test statistics have been reported in Johansen and Juselius (1990).

Granger (1988) points out that if two series are cointegrated, then there must be Granger-causation in at least one direction. Assume that the JJ test procedure indicates that wage and price series are cointegrated and that the estimated cointegrating relationship is

$$P_t = \delta(w - q)_t + U_{1t}, \delta > 0,$$

where U_1 is the random disturbance term. Equation (5) then implies that there exists an error-correction representation of price and wage series of the form

[7] The rank of a matrix is the number of linearly independent columns (or rows) in that matrix.

$$\Delta P_t = a_0 + \sum_{s=1}^{k} a_{1s}\Delta P_{t-s} + \sum_{s=1}^{k} a_{2s}\Delta(w-q)_{t-s} + \lambda_1[P_{t-1} - \delta(w-q)_{t-1}] + \epsilon_{1t} \tag{6.1}$$

$$\Delta(w-q)_t = b_0 + \sum_{s=1}^{k} b_{1s}\Delta(w-q)_{t-s} + \sum_{s=1}^{k} b_{2s}\Delta P_{t-s} + \lambda_2[P_{t-1} - \delta(w-q)_{t-1}] + \epsilon_{2t}, \tag{6.2}$$

where all variables are as defined before and where one of $\lambda_1, \lambda_2 \neq 0$.[8] Equation (6) indicates that whenever the price level P_{t-1} deviates from the long-run value $\delta(w-q)_{t-1}$, then either prices or wages or both adjust so as to keep these two series together in the long run. Lagged levels of the variables now enter the VAR via the error-correction term $P_{t-1} - \delta(w-q)_{t-1}$. Test of the hypothesis that wages do not Granger-cause prices is that all $a_{2s} = 0$ and/or $\lambda_1 = 0$. Hence, the presence of Granger-causality is also examined by testing whether one or both of $\lambda_1, \lambda_2 \neq 0$.

Estimation and Tests of Hypotheses in Cointegrating Vectors

Suppose that the JJ test procedure indicates that price and wage series are cointegrated. In order to examine the nature of long-term correlations between price and wage series, the cointegrating wage and price regressions are estimated using the dynamic OLS procedure described in Stock and Watson (1993).[9] The dynamic versions of these regressions are

$$P_t = a_0 + a_1(w-q)_t + \sum_{s=-k}^{k} a_{2s}\Delta(w-q)_{t-s} + U_{1t} \tag{7.1}$$

$$(w-q)_t = b_0 + b_1 P_t + \sum_{s=-k}^{k} b_{2s}\Delta P_{t-s} + U_{2t}, \tag{7.2}$$

where all variables are as defined before and where U_1 and U_2 are random disturbance terms. Equation (7) includes, in addition, past, current and future values of first differences of the right-hand variables that appear in the cointegrating regression. Since the random disturbance terms, U_1 and U_2, may be

[8] If $\lambda_1 = \lambda_2 = 0$, then the matrix Π in (5) has a zero rank, indicating the absence of any long-run relationship between wage and price series.

[9] The JJ test procedure also provides maximum likelihood estimates of the cointegrating price and wage regressions. These estimates, though superior asymptotically, do not behave well in small samples. In contrast, Stock and Watson's (1993) dynamic OLS behaves well in small samples.

serially correlated, standard test statistics corrected for the presence of serial correlation are used to test hypotheses in (7). Thus, wages are not significantly correlated with the price level in the long run if the hypothesis $a_1 = 0$ or $b_1 = 0$ is not rejected.

Testing for Unit Roots and Mean Stationarity

The cointegration test requires that the time series in X_t be integrated of order one.[10] That is, the data should be stationary in their first differences but not in levels. To determine the order of integration, I use the test procedure suggested by Dickey and Fuller (1979). In particular, the unit root tests are performed by estimating the Augmented Dickey-Fuller regression of the form

$$y_t = a_0 + \rho y_{t-1} + \sum_{s=1}^{k} a_{2s} \Delta y_{t-s} + \epsilon_t, \quad (8)$$

where y_t is the pertinent variable; ϵ the random disturbance term; and k the number of lagged first differences of y_t necessary to make ϵ_t serially uncorrelated. If $\rho = 1$, y_t has a unit root. The null hypothesis $\rho = 1$ is tested using the t-statistic. The lag length (k) used in tests is chosen using the procedure[11] given in Hall (1990), as advocated by Campbell and Perron (1991).

The unit root tests in (8) test the null hypothesis of unit root against the alternative that y_t is mean stationary (the alternative is trend stationary if a linear trend is included in [8]). Recently, some authors including DeJong et al. (1992) have presented evidence that the Dickey-Fuller tests have low power in distinguishing between the null and the alternative. These studies suggest that in trying to decide whether the time series data are stationary or integrated, it would also be useful to perform tests of the null hypothesis of mean stationarity (or trend stationarity). Thus, tests of mean stationarity are performed using the procedure advocated by Kwiatkowski, Phillips, Schmidt, and Shin (1992). The test, hereafter denoted as the KPSS test, is implemented by calculating the test statistic

$$\hat{n}_u = \frac{1}{T^2} \sum_{t=1}^{T} S_t^2 / \hat{\sigma}^2(k),$$

where $S_t = \sum_{i=1}^{t} e_i, t = 1, 2, \ldots T$; e_t is the residual from the regression of y_t on an intercept; S_t is the partial sum of the residuals e; $\hat{\sigma}(k)$ is a consistent

[10] The series is said to be integrated of order one if it is stationary in first differences.

[11] The procedure is to start with some upper bound on k, say $k$max, chosen a priori (eight quarters here). Estimate the regression (8) with k set at $k$max. If the last included lag is significant (using the standard normal asymptotic distribution), select $k = k$max. If not, reduce the order of the estimated autoregression by one until the coefficient on the last included lag (on Δy in [8]) is significant. If none is significant, select $k = 0$.

estimate of the long-run variance[12] of y; and T is the sample size. The statistic $\hat{n}_u$ has a nonstandard distribution and its critical values have been provided by Kwiatkowski et al. (1992). The null hypothesis of mean stationarity is rejected if $\hat{n}_u$ is large. Thus, a time series y_t is considered unit root nonstationary if the null hypothesis that y_t has a unit root is not rejected by the Augmented Dickey-Fuller test and the null hypothesis that it is mean stationary is rejected by the KPSS test.

2. EMPIRICAL RESULTS

This section presents empirical results. In particular, I examine the long- and short-term interactions between wages and prices in a trivariable system consisting of the price level, productivity-adjusted wage, and a demand pressure variable. The price level is measured either by the log of the implicit GDP deflator ($\ln P$) or by the log of the consumer price index (ln CPI); productivity-adjusted wage by the log of the index of unit labor costs of the nonfarm business sector (ln ULC); and demand pressure variable by the log of real over potential GDP (denoted as GAP). Unit labor costs are measured as compensation per hour divided by output per hour. Since supply shocks could have important short-run effects on wages and prices, tests of Granger-causality are conducted including some of these in the trivariable system. The supply shocks considered here include relative prices of energy and imports. Dummy variables for the period of President Nixon's wage and price controls and for the period immediately following the wage and price controls are also included.[13] The data used are quarterly and cover the sample period 1955Q1 to 1992Q4.

Test Results for Unit Roots and Mean Stationarity

In order to determine first whether linear trend is present in the data, Table 1 presents t-statistics on constant and time variables from regressions of the form

[12] The residual e_t is from the regression $y_t = a + b\text{Time} + e_t$. The variance of y_t is the variance of the residuals from this regression and is estimated using the Newey and West's (1987) method as

$$\hat{\sigma}(k) = \frac{1}{T}\sum_{t=1}^{T} e_t^2 + \frac{2}{T}\sum_{s=1}^{T} b(s,k) \sum_{t=s+1}^{T} e_t e_{t-s},$$

where T is the sample size; the weighing function $b(s,k) = 1 + \frac{s}{1+k}$; and k is the lag truncation parameter. The lag parameter was set at $k = 8$. For another simple description of the test procedure, see Ireland (1993).

[13] The relative price of energy is the ratio of the producer price index for fuels, petroleum, and related products to the producer price index for all commodities, and the relative price of imports is the ratio of the implicit deflator for imports to the implicit GNP deflator. The dummy variable for the period of price controls is 1 for 1971Q3 to 1974Q1 and 0 otherwise. The dummy variable for the period immediately following price controls is 1 for 1974Q2 to 1974Q4 and 0 otherwise. The data on prices, unit labor costs, and real GDP are from the Citibase data bank and that on potential GDP from the Board of Governors of the Federal Reserve System.

$$\Delta X_t = a + b\text{Time}_t + \sum_{s=1}^{k} C_s \Delta X_{t-s} + U_t,$$

where X_t is the pertinent variable; U_t a random disturbance term; and k the number of lagged first differences of X_t needed to make U_t serially uncorrelated. If the t-statistic on the constant is large, then X_t has linear trend. In addition, if the t-statistic on the time variable is large, then X_t has quadratic trend. As can be seen, the t-statistics presented in Table 1 are not large for $\ln P$, ln CPI, ln ULC, and GAP, indicating that linear or quadratic trends are not present in any of these time series. Hence, linear trend is not included in tests of unit roots where the alternative hypothesis now is that of mean, not trend, stationarity.

Table 1 Tests for Trends, Unit Roots, and Mean Stationarity

Series X	**Panel A** t-statistics for a Regression of ΔX on:			**Panel B** Tests for Unit Roots				**Panel C** Tests for Mean Stationarity
	Constant	Trend	k	$\hat{\rho}$	$t_{\hat{\rho}}$	k	Confidence Interval for ρ	$\hat{n}_u$
$\ln P$	1.5	.1	3	.99	-0.3	3	(1.00, 1.03)	1.28**
ln CPI	1.2	.1	8	1.00	-0.5	8	(1.00, 1.03)	1.27**
ln ULC	1.5	.3	3	.99	0.0	3	(0.99, 1.03)	1.19**
GAP	-0.4	.2	8	.93	-2.8*	1	(0.83, 1.00)	.23
$\Delta \ln P$				.88	-2.3	2	(0.88, 1.00)	.42
Δ ln CPI				.88	-2.5	8	(0.85, 1.01)	.50
Δ ln ULC				.72	-3.5**	2	(0.77, 0.96)	.35

* Significant at the 10 percent level.
** Significant at the 5 percent level.

Notes: P is the implicit GNP deflator; CPI is the consumer price index; ULC is the unit labor cost; and GAP is the logarithm of real GDP to potential GDP. ln is the natural logarithm and Δ the first-difference operator. The sample period studied is 1955Q1–1992Q4. The t-statistics in Panel A above are from regressions of the form $\Delta X_t = a_0 + a_1\text{TREND} + \sum_{s=1}^{k} a_s \Delta X_{t-s}$, where X is the pertinent series. ρ and t-statistics ($t_{\hat{\rho}}$) for $\rho = 1$ in Panel B above are from the Augmented Dickey-Fuller regression of the form $X_t = a_0 + \rho X_{t-1} + \sum_{s=1}^{k} a_s \Delta X_{t-s}$. The 5 and 10 percent critical values for $t_{\hat{\rho}}$ are -2.9 and -2.6. The number of lagged first differences (k) included in these regressions are chosen using the procedure given in Hall (1990), with maximum lags set at eight quarters. The confidence interval for ρ is constructed using the procedure given in Stock (1991).

The test statistic $\hat{n}_u$ in Panel C above is the statistic that tests the null hypothesis that the pertinent series is mean stationary. The 5 and 10 percent critical values for $\hat{n}_u$ given in Kwiatkowski et al. (1992) are .463 and .574.

Tests for unit roots and mean stationarity are also presented in Table 1. As can be seen, the t-statistic ($t_{\hat{\rho}}$) that tests the null hypothesis that a pertinent time series has a unit root is small for $\ln P$, ln CPI, and ln ULC, but large for GAP. On the other hand, the statistic $\hat{n}_u$ that tests the null hypothesis that a pertinent time series is mean stationary is large for $\ln P$, ln CPI, and ln ULC, but small for GAP. These results indicate that the time series $\ln P$, ln CPI, and ln ULC have a unit root by the ADF test and are not mean stationary by the KPSS test. The GAP variable, on the other hand, does not have a unit root by the ADF test and is mean stationary by the KPSS test. Thus, the wage and price series used here are nonstationary in levels, whereas the demand pressure variable GAP is stationary in levels.

As indicated before, the series has a unit root if $\rho = 1$. Table 1 contains estimates of ρ and their 95 percent confidence intervals.[14] As can be seen, the estimated intervals contain the value $\rho = 1$ and are very tight for $\ln P$, ln CPI, and ln ULC. In contrast, the estimated interval for ρ is fairly wide for the GAP series (.83 to 1.0 for GAP vs. .99 to 1.03 for others). These results further corroborate the evidence above that $\ln P$, ln ULC, and ln CPI each have a unit root whereas the GAP series does not.

The unit root and mean stationary tests using first differences of $\ln P$, ln CPI, and ln ULC are also presented in Table 1. The results here are mixed. The inflation series, $\Delta \ln P$ and Δ ln CPI, have a unit root by the ADF test but are mean stationary by the KPSS test. The 95 percent confidence interval for ρ is (.88, 1.0) for $\Delta \ln P$ and (.85, 1.0) for Δ ln CPI. These confidence intervals are quite wide, indicating that ρ could as well be below unity (say, $\rho = .8$) and thus the inflation series could as well be stationary. The wage growth series Δ ln ULC, on the other hand, does not have a unit root by the ADF test and is mean stationary by the KPSS test. These results indicate that the wage growth series is mean stationary. The empirical work presented hereafter also treats the inflation series as mean stationary.[15]

Cointegration Test Results

The results presented in the previous section indicate that the price and wage series used here have stochastic, not deterministic, trends. I now examine

[14] The confidence interval for ρ is constructed using the procedure given in Stock (1991). The 95 percent confidence interval provides the range which may contain the true value of ρ with some probability (.95).

[15] These results differ somewhat from those given in Mehra (1991). The unit root tests given in Mehra (1991) were performed including a linear trend and indicated that gap, inflation ($\Delta \ln P$), and growth in unit labor costs (Δ ln ULC) series are nonstationary. The additional test results—such as those on linear trend, mean stationarity, and the confidence intervals for ρ—presented here, however, indicate that inflation, the output gap, and growth in unit labor costs could as well be stationary.

Table 2 Cointegration Test Results

System	k^a	Trace Test	Maximum Eigenvalue Test
($\ln P$, ln ULC)	3	17.2*	17.2*
(ln CPI, ln ULC)	5	20.2**	19.8**

[a] The lag length k was selected using the likelihood ratio test procedure described in footnote 16 of the text.
* Significant at the 10 percent level.
** Significant at the 5 percent level.

Notes: Trace and maximum eigenvalue tests are tests of the null hypothesis that there is no cointegrating vector in the system. The 5 percent and 10 percent critical values are 17.8 and 15.6 for the Trace statistic and 14.6 and 12.8 for the maximum eigenvalue statistic. Critical values are from Johansen and Juselius (1990).

whether there exists a long-run equilibrium relationship between $\ln P$ and ln ULC or between ln CPI and ln ULC, using the test of cointegration.

Table 2 presents cointegration test results using the JJ procedure.[16] As can be seen, trace and maximum eigenvalue test statistics that test the null that there is no cointegrating vector are large and significant, indicating that the wage and price series are cointegrated. The cointegrating price and wage regressions estimated using the dynamic OLS procedure are reported in Table 3. χ_1^2 is the Chi-square statistic that tests the null hypothesis that the coefficient on ln ULC in the price regression is zero. Similarly, χ_2^2 tests the null that the coefficient on $\ln P$ (or on ln CPI) is zero in the wage regression. As can be seen, χ_1^2 and χ_2^2 take large values and are significant, indicating that prices and wages are significantly correlated in the long run.[17] Furthermore, the estimated coefficients that appear on price and wage variables in these cointegrating regressions are positive and not far from unity. This indicates that wage and price series may grow at similar rates in the long run.

[16] The lag length parameter (k) for the VAR model was chosen using the likelihood ratio test described in Sims (1980). In particular, the VAR model initially was estimated with k set equal to a maximum of eight quarters. This unrestricted model was then tested against a restricted model where k is reduced by one, using the likelihood ratio test. The lag length finally selected in performing the JJ procedure is the one when the restricted model is rejected.

[17] The relevant statistics have a Chi-square, not an F, distribution because standard errors have been corrected for the presence of moving average serial correlation. The order of the moving average correction was determined by examining the autocorrelation of the residuals at various lags.

Table 3 Cointegrating Vectors; Dynamic OLS

Price Regressions	Wage Regressions
$\ln P_t = -.29 + 1.03 \ln \mathrm{ULC}_t$ $\chi_1^2 = 41.3$	$\ln \mathrm{ULC}_t = .31 + .96 \ln P_t$ $\chi_2^2 = 83.9$
$\ln \mathrm{CPI}_t = -.23 + 1.05 \ln \mathrm{ULC}_t$ $\chi_1^2 = 111.1$	$\ln \mathrm{ULC}_t = .23 + .94 \ln \mathrm{CPI}_t$ $\chi_2^2 = 196.8$

Notes: All regressions are estimated by the dynamic OLS procedure given in Stock and Watson (1993), using eight leads and lags of first differences of the relevant right-hand side explanatory variables. χ_1^2 is the Chi-square statistic that tests the null hypothesis that ln ULC is not significant, whereas χ_2^2 tests the null that $\ln P$ or ln CPI is not significant. Both statistics are distributed Chi-square with one degree of freedom. The standard errors in these regressions were corrected for the presence of moving average serial correlation.

Granger-Causality Test Results

The presence and nature of short-term interactions between wage and price series are investigated by estimating regressions of the form

$$\Delta \ln P_t = a_0 + \sum_{s=1}^{k1} a_{1s} \Delta \ln P_{t-s} + \sum_{s=1}^{k2} a_{2s} \Delta \ln \mathrm{ULC}_{t-s} + \sum_{s=1}^{k3} a_{3s} \mathrm{GAP}_{t-s} + \lambda_1 \hat{U}p_{t-1} + \epsilon_{1t} \tag{9}$$

$$\Delta \ln \mathrm{ULC}_t = b_0 + \sum_{s=1}^{k1} b_{1s} \Delta \ln \mathrm{ULC}_{t-s} + \sum_{s=1}^{k2} b_{2s} \Delta \ln P_{t-s} + \sum_{s=1}^{k3} b_{3s} \mathrm{GAP}_{t-s} + \lambda_2 \hat{U}w_{t-1} + \epsilon_{2t}, \tag{10}$$

where P is the price level measured either by the implicit GDP deflator or by the consumer price index; $\hat{U}p$ the residual from the cointegrating price regression; $\hat{U}w$ the residual from the cointegrating wage regression;[18] and $k1$, $k2$, and $k3$ the lag lengths on various variables needed to make random disturbances (ϵ_1, ϵ_2) serially uncorrelated. Wages do not Granger-cause prices if all $a_{2s} = 0$ and/or $\lambda_1 = 0$, and prices do not Granger-cause wages if all $b_{2s} = 0$ and/or $\lambda_2 = 0$.

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[18] The appearance of error-correction terms in (9) and (10) follows directly from equation (5). If wage and price series are cointegrated, then the term ΠX_{t-k} in equation (5) captures coefficients that appear on the linear combination of wage and price variables. That is also demonstrated in equations (6.1) and (6.2).

Table 4 Error-Correction Coefficients and F Statistics for Granger-Causality; Implicit GDP Deflator

	Statistics from Price Regressions				Statistics from Wage Regressions			
Sample Period	**Lag Lengths ($k1$, $k2$, $k3$)**	**$\lambda 1$ (t-value)**	**F1**	**d.f.**	**$\lambda 2$ (t-value)**	**F2**	**d.f.**	**χ^2_w**
1956Q1–	(4,0,0)	.02(1.6)			−.14(3.9)			
1992Q4	(8,0,0)	.03(1.9)			−.17(4.5)			5.8
	(4,4,4)	−.01(0.4)	0.3	4,116	.05(1.1)	7.1**	4,124	
	(8,8,8)	−.01(0.2)	1.6	8,104	.06(1.2)	1.9*	8,112	
	(7,8,2)[a]	−.00(0.1)	1.2	8,111				
	(0,4,1)[a]				−.02(0.5)	22.4**	4,131	
1956Q1–	(4,0,0)	.03(1.5)			−.17(3.9)			
1979Q3	(8,0,0)	.04(1.7)			−.18(4.1)			7.0
	(4,4,4)	.04(0.9)	0.1	4,63	.06(1.0)	4.8**	4,71	
	(8,8,8)	.05(0.9)	0.8	8,51	.14(1.3)	0.8	8,59	
	(7,0,0)[a]	.03(1.6)						
	(4,4,1)[a]				.03(0.5)	7.3**	4,74	

[a] Lag lengths chosen using the procedure given in Hall (1990).
* Significant at the 10 percent level.
** Significant at the 5 percent level.

Notes: The price regressions are of the form

$$\Delta \ln P_t = a_0 + \sum_{s=1}^{k1} a_{1s} \Delta \ln P_{t-s} + \sum_{s=1}^{k2} a_{2s} \Delta \ln \mathrm{ULC}_{t-s} + \sum_{s=1}^{k3} a_{3s} \mathrm{GAP}_{t-s} + \lambda_1 U_{1t-1}$$

and wage regressions are of the form

$$\Delta \ln \mathrm{ULC}_t = b_0 + \sum_{s=1}^{k1} b_{1s} \Delta \ln \mathrm{ULC}_{t-s} + \sum_{s=1}^{k2} b_{2s} \Delta \ln P_{t-s} + \sum_{s=1}^{k3} b_{3s} \mathrm{GAP}_{t-s} + \lambda_2 U_{2t-1}.$$

U_1 is the residual from the cointegrating price regression and U_2 from the cointegrating wage regression, both reported in Table 3. F1 tests all $a_{2s} = 0$, F2 tests all $b_{2s} = 0$, and d.f. is the degree-of-freedom parameter for the F statistic given in the relevant row. The price regressions also included eight past values of the relative prices of energy and imports and dummies for President Nixon's price controls. The wage regressions included eight past values of the relative price of imports and price control dummies. χ^2_w is the Lagrange multiplier test for the hypothesis that eight lags of the relative price of energy do not enter the wage regression (the 5 percent critical value is 15.5).

Tables 4 and 5 report estimates of λ_1 and λ_2 (with t-statistics in parentheses) from regressions of the form (9) and (10). In Table 4 the price series used is the implicit GDP deflator and in Table 5 it is the consumer price index. The regressions are estimated using some arbitrarily chosen lag lengths ($k1$, $k2$, $k3$) as well as those chosen on the basis of the procedure given in Hall (1990). In addition, the results are presented for the subperiod 1956Q1 to 1979Q3. The price regression (9) estimated here also included eight past values of relative

Table 5 Error-Correction Coefficients and F Statistics for Granger-Causality; Consumer Price Index

		Statistics from Price Regressions			Statistics from Wage Regressions			
Sample Period	**Lag Lengths ($k1$, $k2$, $k3$)**	**$\lambda 1$ (t-value)**	**F1**	**d.f.**	**$\lambda 2$ (t-value)**	**F2**	**d.f.**	χ^2_w
1956Q1–1992Q4	(4,0,0)	−.05(2.2)			−.10(1.8)			
	(8,0,0)	−.05(2.2)			−.12(2.3)			18.1
	(4,4,4)	−.05(2.1)	0.70	4,116	−.04(0.8)	4.4**	4,132	
	(8,8,8)	−.02(0.6)	1.58	8,104	−.04(0.8)	0.5	8,120	
	(8,0,2)[a]	−.05(2.7)		8,104				
	(0,1,1)[a]				−.00(0.4)	68.7**	1,142	
1956Q1–1979Q3	(4,0,0)	−.05(1.4)			−.01(0.4)			
	(8,0,0)	−.04(1.4)			−.24(2.5)			
	(4,4,4)	−.10(2.4)	1.4	4,63	−.08(1.0)	4.9**	4,63	
	(8,8,8)	.01(0.2)	1.5	8,51	−.15(1.1)	1.1	8,51	
	(8,4,1)[a]	.11(2.9)	2.2*	4,62				
	(0,1,1)[a]				−.05(0.8)	14.5**	1,73	

Notes: See notes in Table 4. The regressions are estimated using the consumer price index.

prices of energy and imports, and "on" and "off" dummies for President Nixon's price controls. The wage regression (10) included, in addition, eight past values of the relative price of imports and price control dummies. Coefficients for the relative price of energy were not significant in such regressions.[19]

If we focus on results for the general price level presented in Table 4, it is clear that λ_1 is generally not statistically significant whereas λ_2 is significant (see t-statistics in Table 4). Moreover, other lags of $\Delta \ln$ ULC when included in price regressions are not statistically significant, whereas other lags of $\Delta \ln P$ when included in wage regressions are statistically significant (compare F1 and F2 statistics in Table 4). These results are consistent with the presence of Granger-causality, not from wages to prices, but from prices to wages.

The results using consumer prices are somewhat different from those using the general price level. As can be seen from Table 5, λ_1 is generally statistically significant, even though other lags of $\Delta \ln$ ULC when included in price regressions are not (see t- and F1 statistics in Table 5). On the other hand, λ_2

[19] Whether or not some of these supply shocks enter price and wage regressions was first tested using the Lagrange-multiplier (LM) test for omitted variables (Engle 1984). An LM test for k omitted variables is constructed by regressing the equation's residuals on both the original regressors and on the set of omitted variables. If the omitted variables do not belong in the equation, then multiplying the R^2 statistic from this regression by the number of observations will produce a statistic asymptotically distributed χ^2 with k degrees of freedom.

or other lags of $\Delta \ln$ CPI when included in wage regressions are statistically significant (see t- and F2 statistics in Table 5). These results are consistent with the presence of Granger-causality between prices and wages with feedbacks in both directions.

3. CONCLUDING OBSERVATIONS

A central proposition in the expectations-augmented Phillips-curve model of the inflation process is that prices are marked up over productivity-adjusted labor costs. If that proposition is correct, then long-run movements in prices and labor costs must be correlated. Moreover, we should find that short-run movements in labor costs help predict short-run movements in the price level. The evidence reported here indicates that these implications are consistent with the data when prices are narrowly measured by the consumer price index but not when they are broadly defined by the implicit GDP deflator. For the latter measure, short-run movements in labor costs have no predictive content for the future price level. The general price level and unit labor costs are still correlated in the long run. But the presence of this correlation appears to be due to Granger-causality running from the general price level to labor costs, not the other way around. Huh and Trehan (1992) report similar results using business sector price and wage data.

The finding that consumer prices and unit labor costs are Granger-causal with feedbacks in both directions differs from the one in Barth and Bennett (1975), which found Granger-causality running one way from consumer prices to wages. The empirical work in Barth and Bennett, however, does not test for the presence of Granger-causality occurring via the error-correction term. If we were to ignore this channel, other test results presented here are also consistent with the presence of Granger-causality running one way from consumer prices to wages (compare F1 and F2 statistics in Table 5).

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