# Do Wages Help Predict Inflation?

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*The inclusion of unit labor costs in forecasts of consumer price inflation provides no significant improvement in forecasting errors, especially in recent years.*

Gauging inflationary pressures is a perennial concern for monetary policymakers and financial market participants. For example, during 1994, the Federal Reserve tightened monetary policy in response to concerns about building inflationary pressures and higher inflation in the future. Many people attributed the pursuant 1995 slowdown in economic activity largely to these Federal Reserve actions.

Unfortunately, the performance of many once reliable guides of future inflation, such as the growth of monetary aggregates, has deteriorated in recent years. This deterioration has resulted from, among other things, financial innovations that have changed the relationships between financial variables and economic activity. The deterioration of once reliable inflation guides has led policymakers and financial markets to monitor a broad range of inflation indicators. Because labor costs make up more than two-thirds of the total cost of producing goods and services in the United States, one of these indicators is unit labor costs, or wages adjusted for changes in labor productivity. Indeed, many analysts currently cite the lack of accelerating unit labor costs as grounds for believing that inflation will not increase any time soon.

Research on the relationship between unit labor costs and inflation has focused on whether higher labor costs precede higher inflation, or vice versa. In statistical jargon, the research has focused on whether labor costs Granger-cause inflation.[1] Recent research by Mehra (1993, 1991) that utilizes newly developed statistical techniques yields mixed results. Mehra (1993) finds that when consumer prices serve as the measure of prices, unit labor costs and prices are correlated in the long run. The study also finds that this correlation is present because Granger causality is running in both directions,[2] which implies that unit labor costs contain information about future consumer prices. However, Mehra (1991) finds that when the gross domestic product (GDP) deflator is used as the measure of prices, a long-run correlation still exists, but its source is Granger causality that runs only from prices to wages. Therefore, in this case, unit labor costs have no information content for future movements in prices.

The purpose of this article is twofold. The first is to examine how much forecasting power unit labor costs have for future consumer prices. Is the attention paid to unit labor costs as an inflation indicator justified? While Mehra finds that unit labor costs Granger-cause consumer prices, he does not examine the extent of unit

Figure 1
**Growth of Unit Labor Costs and Consumer Price Inflation, 1957–93**


labor costs' predictive power for out-of-sample forecasts of inflation.[3] The second purpose of this study is to examine whether the relationship between unit labor costs and consumer prices is stable over time.[4] As with the inflation-indicator properties of the monetary aggregates, have the indicator properties of unit labor costs deteriorated in recent years?

Our empirical strategy is to first take a preliminary look at the raw data and the data transformed by a filter designed by Baxter and King (1995). Next, we carry out Granger causality tests and a stability analysis of those tests. Finally, we examine the forecasting ability of unit labor costs for consumer price inflation (CPI). Our main finding is that the inclusion of unit labor costs in forecasts of consumer price inflation provides no significant improvement in forecasting errors, especially in recent years.

## A preliminary look at labor costs and prices

Figure 1 plots year-over-year growth of unit labor costs and consumer price inflation, excluding food and energy (CPIC).[5] The high correlation between movements in labor costs and inflation demonstrates why analysts have paid close attention to labor costs when assessing inflation. However, what is not clear from the figure is whether movements in labor costs precede movements in inflation, or vice versa. In other words, it is not clear from Figure 1 whether movements in labor costs help to forecast future movements in inflation.

Notice also from Figure 1 that there appears to be a potential break in the relationship between labor costs and inflation sometime during the early 1980s. The growth of labor costs seems to be persistently lower than inflation growth during the 1980s, and the contemporaneous correlation between the two variables appears lower.

Using a filter methodology developed by Baxter and King (1995), we can divide labor cost growth and inflation into their long-run and business-cycle components. The results of doing this are shown in Figures 2A and 2B and illustrate that labor cost growth and inflation are correlated at both the business-cycle frequency and in their trend, or long-run, movements. Table 1 provides correlations from the raw data and for the trend and cycle components for the entire sample and for two subsamples.[6] The correlations for wages leading prices at the trend and business-cycle frequencies (negative *k*s) are positive, although higher at the trend frequency, supporting the view that movements in wages could help predict future movements in prices. Additionally, these correlations seem to be con-

Figure 2A
**Inflation and Wage Growth Components**

*Trend components*

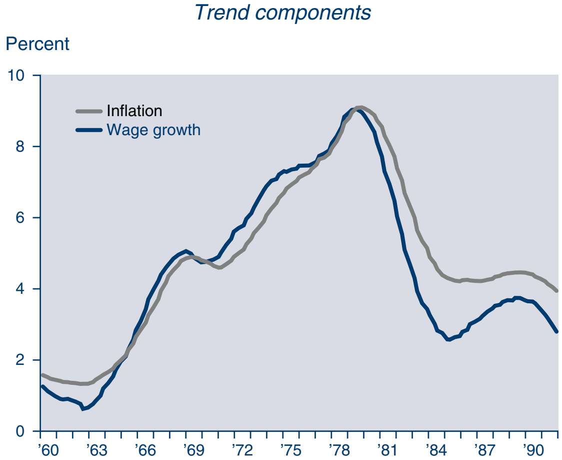



Figure 2B
**Inflation and Wage Growth Components**

*Cycle components*

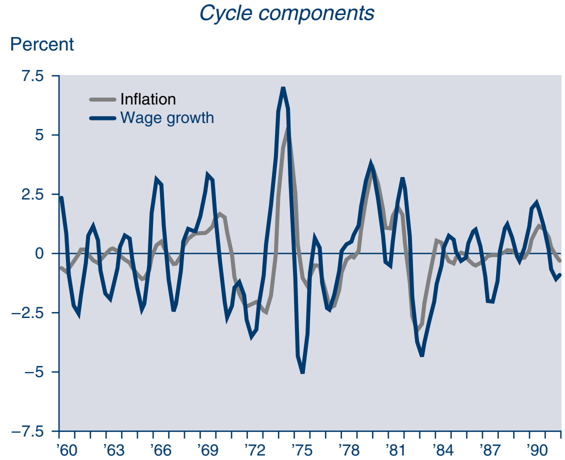


Table 1
**Cross Correlation: Inflation and Wage Growth**

| k | 1965:4–80:4 | 1981:1–94:4 | 1957:1–94:4 |
|---|---|---|---|
| Raw data | | | |
| −4 | .37 | .05 | .54 |
| −3 | .50 | .19 | .60 |
| −2 | .59 | .44 | .68 |
| −1 | .56 | .45 | .67 |
| 0 | .60 | .51 | .67 |
| 1 | .38 | .60 | .63 |
| 2 | .19 | .51 | .54 |
| 3 | .04 | .27 | .42 |
| 4 | −.09 | .10 | .35 |
| Trend components | | | |
| −4 | .79 | .58 | .94 |
| −3 | .85 | .68 | .95 |
| −2 | .90 | .77 | .96 |
| −1 | .94 | .87 | .96 |
| 0 | .98 | .96 | .96 |
| 1 | .90 | .82 | .93 |
| 2 | .83 | .68 | .90 |
| 3 | .75 | .54 | .87 |
| 4 | .68 | .41 | .83 |
| Cycle components | | | |
| −4 | .45 | −.35 | .31 |
| −3 | .66 | −.20 | .49 |
| −2 | .81 | .10 | .64 |
| −1 | .83 | .44 | .72 |
| 0 | .68 | .72 | .65 |
| 1 | .38 | .84 | .46 |
| 2 | .03 | .76 | .18 |
| 3 | −.25 | .51 | −.09 |
| 4 | −.40 | .20 | −.27 |

NOTES: The reported coefficients show the correlation between inflation at time $t$ and wage growth at time $t + k$.

sistent with a potential breakpoint sometime in the early 1980s: for the raw data and both filter components, the wage leading inflation coefficients (negative-signed $k$s) drop in the 1980s. Additionally, for the raw data and the cycle data, the inflation leading wages coefficients (positive signed $k$s) increase during the 1980s.

## Granger-causality results

**Whole sample.** We use both consumer prices for all items (CPI) and consumer prices excluding food and energy (CPIC) as our price measures.[7] Unit labor costs are for the nonfarm business sector. As a preliminary step to the formal causality tests, we have to determine the stationarity characteristics of the time series.[8] We choose the augmented Dickey–Fuller method (ADF) to conduct the tests for each variable in levels, first differences, and second differences. Table 2 summarizes the results and shows that unit labor costs and both price measures are integrated of order two, denoted by I(2).[9]

Granger causality tests with the variables in second differences will still be misspecified if inflation growth and wage growth are cointegrated and converge to a stationary long-run equilibrium relationship.[10] If the series are cointegrated, an error-correction term must be included in the causality test. This necessity follows from Engel and Granger's (1987) findings that if two variables are cointegrated, an error-correction model for the variables is present and that not including the error-correction term can lead to faulty inferences. Furthermore, cointegration between two variables implies Granger causation in at least one direction. The presence of cointegration provides a dynamic framework in which an error-correction term represents deviations from a long-run cointegrating relationship, while lagged difference terms represent short-run dynamics.

To estimate the possibility of a cointegrating relationship between the first difference of prices and unit labor costs, we use the Dynamic OLS (DOLS) procedure of Stock and Watson (1993). This procedure entails regressing one of the I(1) variables on the other I(1) variable, and lags and leads of the first differences of the I(1) variables. With standard errors corrected for serial correlation, one can make valid inferences from each coefficient estimate. The procedure is described by the following equations:

$$(1) \qquad \Delta p_t = \alpha_p + \beta_p \Delta w_t + \sum_{i=-k}^{k} \gamma_{pi} \Delta^2 w_{t-i} + \epsilon_{pt}$$

and

$$(2) \qquad \Delta w_t = \alpha_w + \beta_w \Delta p_t + \sum_{i=-k}^{k} \gamma_{wi} \Delta^2 p_{t-i} + \epsilon_{wt},$$

where $p$ and $w$ are the logarithms of prices and unit labor costs and $\Delta$ is the difference operator. Table 3 shows the results of testing the $\alpha$'s and $\beta$'s.[11] Both $\beta_p$ and $\Delta_w$ are significant at the 1-percent level, but the $\alpha$'s are only significant with CPIC. However, the augmented Dickey–Fuller tests for the cointegrating residuals confirm a stationary relationship between the growth of both price measures and the growth of unit labor costs, implying cointegration.[12]

We are now ready to conduct the Granger causality tests. To examine the causal relationship between inflation and wage growth, we estimate the following bivariate models:

$$(3) \qquad \Delta^2 p_t = a_p + b_p \left( \Delta p - \alpha_p - \beta_p \Delta w \right)_{t-1} + \sum_{i=1}^{k} c_{pi} \Delta^2 p_{t-i} + \sum_{i=1}^{k} d_{pi} \Delta^2 w_{t-i} + \epsilon_{pt}$$

and

$$\Delta^2 w_t = a_w + b_w\left(\Delta w - \alpha_w - \beta_w \Delta p\right)_{t-1} + \sum_{i=1}^{k} c_{wi}\Delta^2 p_{t-i} + \sum_{i=1}^{k} d_{wi}\Delta^2 w_{t-i} + \epsilon_{wt}, \tag{4}$$

where the terms in parentheses are the error-correction terms and are estimated by DOLS. The hypothesis of no causality from wage to inflation is rejected if $b_p$ and/or all $d_p$'s are significantly different from zero.

Our results for the causality tests are summarized in Table 4.[13] Whether wage growth Granger-causes inflation depends on the choice of the price series. For CPIC, wage growth is significant at the 1-percent level, implying causality. However, for CPI, wage growth is not significant, implying no causality.[14] The results also show that inflation always Granger-causes wage growth, regardless of the choice of the price series. It is noteworthy that the error-correction terms play a crucial role for the rejection of the no-causality hypotheses.

**Granger-causality tests: A stability analysis.** There are two potential sources of instability in the Granger causality test results presented above. First, there may be instability in the cointegrating relationship. Second, there may be instability in the short-run dynamics or in the Granger regressions themselves.

To examine the stability of the cointegrating relationship, we use Stock and Watson's (1993) formal test for the null hypothesis of a constant cointegrating relationship against the alternative of different cointegrating vectors over various samples. In their test for structural stability, the constant term in the error-correction term remains fixed. In contrast, our alternative test is based on the following regression:

$$\Delta p_t = \alpha_p + \beta_p \Delta w_t + \left(\theta_p + \delta_p \Delta w_t\right)1\left(t > \tau\right) + \sum_{i=-k}^{k} \gamma_{pi}\Delta^2 w_{t-i} + \epsilon_{pt}, \tag{5}$$

where $\tau$ is the possible break date and $1(t > \tau)$ is the indicator function equal to 1 if $t > \tau$, 0 otherwise. The joint significance of $\theta_p$ and $\delta_p$ implies rejection of the null hypothesis of a stable long-run relationship. We conduct the joint test for $\theta_p$ and $\delta_p$ sequentially with all possible $\tau$'s, which covers 70 percent of the sample period. For each $\tau$, we get a $\chi^2$ statistic indicating the significance of $\theta_p$ and $\delta_p$. The maximum of the sequence of $\chi^2$ statistics yields a possible break in the cointegrating relationship.[15] It should be noted that the conventional $\chi^2$ critical values are invalid because the timing of the structural change is not specified under the alternative. Hansen (1992) derives a SupF test for parameter instability in the context of cointegrated regression models. The far right column in Table 3 displays the SupF statistics and the selected break dates. Based on Hansen's asymptotic critical values, we discover a significant shift in regime. Depending on the price series, the break dates

Table 2

## Augmented Dickey–Fuller Test Results

| | $\tau$-statistics | Lag order ($k$) | Ljung–Box Q-statistics |
|---|---|---|---|
| Levels | | | |
| CPI core | −2.68 | 3 | 15.70* |
| CPI | −2.76 | 4 | 13.81 |
| Unit labor cost | −2.11 | 3 | 11.08 |
| First-differenced | | | |
| CPI core | −2.44 | 7 | 14.89 |
| CPI | −2.50 | 8 | 11.15 |
| Unit labor cost | −2.34 | 8 | 12.32 |
| Second-differenced | | | |
| CPI core | −4.62*** | 6 | 17.86* |
| CPI | −5.03*** | 8 | 13.69 |
| Unit labor cost | −6.03*** | 8 | 13.35 |

*** = Significance at the 1-percent level.
** = Significance at the 5-percent level.
* = Significance at the 10-percent level.

NOTES: The testing equations are of the form:

$$y_t = \alpha + (\theta t) + \rho y_{t-1} + \sum_{i=1}^{k} \Delta y_{t-i} + \epsilon_t,$$

where the lag length $k$ is determined by the Schwartz information criterion for $1 \le k \le 8$. Ljung–Box Q-statistics are used to check the serial correlation of the residuals. Q(9) and Q(10) are reported for the levels and for the first and second differences, respectively. All variables are in natural logs. The variables in levels are tested for trend stationarity, and the first- and second-differenced variables are tested for difference stationarity.

Table 3

## Dynamic OLS Cointegration Test: Inflation and Wage Growth

| | $\alpha_p$ | $\beta_p$ | Dickey–Fuller $\tau$-statistics | Break date ($\chi^2_2$) |
|---|---|---|---|---|
| CPI | .16 | 1.03 | −4.95*** | 1980:2 (15.97**) |
| | (.84) | (257.1***) | | |
| CPI core | .65 | .92 | −4.04*** | 1980:4 (3.45***) |
| | (14.32***) | (257.2***) | | |

*** = Significance at the 1-percent level.
** = Significance at the 5-percent level.
* = Significance at the 10-percent level.

NOTES: (1) In the test of the significance of $\alpha$ and $\beta$, the reported $\chi^2_1$ statistics use Newey and West (1987) robust standard errors with a truncation lag of 4. Augmented Dickey–Fuller tests are implemented on the cointegrating residual, $\Delta p - \alpha_p - \beta_p \Delta w$, with the lag length of eight determined by Schwartz information criterion.
(2) To search for the significant break dates, $\chi^2_2$ is calculated to test for the changes in a and b of the cointegrating residuals over the 70 percent of the whole sample. The date is chosen based on the largest $\chi^2$ statistic. Hansen's (1992) critical values are 16.2, 12.4, and 1.6 at the 1-, 5-, and 10-percent level, respectively.

Table 4
## Causality Tests for Bivariate ECM

$$\Delta^2 p_t = a_p + b_p\left(\Delta p - \alpha_p - \beta_p \Delta w\right)_{t-1} + \sum_{i=1}^{k} c_{pi}\Delta^2 p_{t-i} + \sum_{i=1}^{k} d_{pi}\Delta^2 w_{t-i} + \epsilon_{pt}$$

$$\Delta^2 w_t = a_w + b_w\left(\Delta w - \alpha_w - \beta_w \Delta p\right)_{t-1} + \sum_{i=1}^{k} c_{wi}\Delta^2 p_{t-i} + \sum_{i=1}^{k} d_{wi}\Delta^2 w_{t-i} + \epsilon_{wt}$$

**Price equation**

| **Null hypothesis** | $b_p = 0$ | $d_{pi} = 0$ ($i$ = 1 ,2,3,4) | $b_p = 0$ and $d_{pi} = 0$ ($i$ = 1,2,3,4) |
|---|---|---|---|
| (p = CPI core) | | | |
| 1957:1–94:4 | 15.5 (.000)*** | 4.23 (.003)*** | 5.23 (.000)*** |
| 1957:1–80:4 | 14.0 (.000)*** | 2.99 (.024)** | 3.79 (.004)** |
| 1981:1–94:4 | .27 (.609) | 1.29 (.289) | 1.04 (.407) |
| (p = CPI) | | | |
| 1957:1–94:4 | 1.80 (.182) | 1.42 (.230) | 1.57 (.172) |
| 1957:1–80:2 | .07 (.796) | .97 (.428) | .81 (.546) |
| 1980:3–94:4 | .59 (.446) | .72 (.585) | .65 (.663) |

**Wage equation**

| **Null hypothesis** | $b_w = 0$ | $c_{wi} = 0$ ($i$ = 1,2,3,4) | $b_w = 0$ and $c_{wi} = 0$ ($i$ = 1,2,3,4) |
|---|---|---|---|
| (p = CPI Core) | | | |
| 1957:1–94:4 | 13.8 (.000)*** | 1.30 (.274) | 6.41 (.000)*** |
| 1957:1–80:4 | 5.56 (.021)** | 1.86 (.125) | 4.46 (.001)*** |
| 1981:1–94:4 | 1.3 (.002)*** | 2.56 (.051)* | 7.15 (.000)*** |
| (p = CPI) | | | |
| 1957:1–94:4 | 33.7 (.000)*** | .82 (.516) | 1.6 (.000)*** |
| 1957:1–80:2 | 28.9 (.000)*** | .69 (.603) | 9.35 (.000)*** |
| 1980:3–94:4 | 12.7 (.000)*** | 1.02 (.409) | 3.41 (.010)** |

*** = Significance at the 1-percent level.
** = Significance at the 5-percent level.
* = Significance at the 10-percent level.

NOTES: The error-correction term in both equations are estimated by Stock and Watson's Dynamic OLS with leads and lags equal to eight. *F*-statistics for the Granger causality tests are reported together with their *p*-values in the parentheses.

are 1980:4 and 1980:2. Therefore, the potential break point apparent in Figure 1 is supported by the formal stability tests.

In sum, the data suggest instability in the cointegrating regression. We now turn to the question of whether this instability affects the nature of the short-run dynamics between changes in wage growth and changes in inflation. In other words, does this instability affect the Granger causality tests? For this purpose, we estimate three Granger causality specifications over the two subsamples and the entire sample.[16] Two specifications are error-correction models (ECM). In the first ECM, the error-correction term is estimated from equations 1 and 2. In the second ECM model, the estimated cointegrating regression is estimated allowing a change in both the constant and the slope across the two subsamples. The third model does not contain an error-correction term.[17] The bottom line is that the Granger causality results from all three of these models turn out to be the same. However, because the original specification (the first ECM model) has a better fit across all samples, both in terms of $R^2$ and in superior forecasts that follow, we report in Table 4 results only from this model. Consequently, for the whole sample, the results of the Granger causality tests are unaffected.

Again, the whole sample results were that unit labor costs Granger-cause CPIC, but not CPI. Interestingly, the results for the subsamples in Table 4 indicate that wage growth Granger-causes CPIC in the pre-1980 sample only, not in the post-1980 sample. In neither subsample do wages cause CPI.[18] However, across all samples and for both price measures, prices cause wages.[19]

Summarizing, the in-sample causality tests indicate that changes in wage growth have information content for changes in core consumer price inflation, but not consumer price inflation. However, the information content for changes in CPIC appears to disappear after 1980. The information content of changes in both measures of inflation for future changes in wages appears to be much more robust. We now turn to an examination of the extent of the information content that wages have for inflation in out-of-sample forecasting exercises.

## Forecasting exercises

**Out-of-sample forecasts.** The forecasting exercises consist of horse races between autoregressive univariate forecasts of inflation and forecasts obtained from the ECM models, which include unit labor costs.[20] The objective is to examine the reduction in forecast errors obtained by including the information content of wages. We carry out forecasts for the level of inflation and wages at forecast horizons of one, four, and eight quarters and for the three samples examined above.

The out-of-sample forecasts provide the real test of how forecasters would have done in real time using productivity-adjusted wages to help predict inflation. The first type of forecast we conduct is to use the ECM model estimated for the 1958–89 sample and then ask how it does in helping predict inflation from 1990 through 1994. As with the in-sample forecasts, we compare the root mean square errors (RMSEs) of the ECM model with a univariate autoregressive model. Each quarter, we update the parameter estimates of both models as the forecasts proceed through the 1990s.

The results are shown in the top section of

Table 5. For CPI inflation, the use of the ECM model actually results in a higher RMSE than the univariate model at all forecast horizons. For CPIC inflation, the use of the ECM model results in modest reductions in RMSEs, particularly at the four-quarter forecast horizon. Note that neither CPI nor CPIC inflation helps forecast wages, as the RMSEs in five out of the six forecasts actually increase with the ECM model.

Because the evidence suggests that there may have been a break in the wage–inflation relationship in the early 1980s, we also conduct out-of-sample forecasts in which the models are estimated using data only from the post-1980 period. In these forecasts, we initially estimate the models over the 1981–89 period and then conduct forecasts for the 1990–94 period. Again, the parameter estimates are updated as the forecasts progress through the 1990s. The results are shown in the middle section of Table 5. They indicate that breaking up the sample does not result in an improvement for the ECM model. In fact, the ECM model does worse.[21] Notice, however, that the RMSEs from the univariate models are the smallest of all the models considered. Therefore, the results indicate that for forecasting CPI and CPIC inflation during the 1990s, the use of only the post-1980 data results in lower forecast errors. Additionally, the inclusion of wage growth actually results in larger errors. Figure 3 plots the forecast errors from both models. Finally, the forecasts of wages indicate that the inclusion of CPIC inflation reduces the forecast errors at four- and eight-quarter horizons but the inclusion of CPI inflation does not result in lower forecast errors.

As a final exercise, we look to see whether, during the late 1970s, wages helped forecast inflation. In these forecasts, we estimate the models using data from the 1958–77 sample and then conduct out-of-sample forecasts for the period 1978–81. The results are shown in the bottom section of Table 5. For CPI, the use of the ECM model does not result in improved forecast errors. For CPIC, the use of the ECM model does result in an improvement, especially at the one- and four-quarter forecast horizons. At the four-quarter horizon, the use of the ECM model results in a 13-percent reduction in RMSE. Figure 4 plots the forecast errors. The results for inflation as a predictor of wage growth are mixed.

Table 5
**Out-of-Sample Root Mean Square Forecast Errors**

| | Inflation | | | Wage growth | | |
|---|---|---|---|---|---|---|
| **Forecast horizons** | **ECM** | **AR(4)** | **(Percent change)** | **ECM** | **AR(4)** | **(Percent change)** |
| Initial in-sample estimation: 1957:1–89:4 | | | | | | |
| Forecasting periods: 1990:1–94:4 | | | | | | |
| (CPIC) | | | | | | |
| 1 | .800 | .808 | (1.10) | 2.000 | 2.082 | (3.96) |
| 4 | .908 | 1.015 | (10.56) | 2.445 | 2.411 | (–1.37) |
| 8 | 1.442 | 1.563 | (7.76) | 3.034 | 2.947 | (–2.97) |
| (CPI) | | | | | | |
| 1 | 1.337 | 1.293 | (–3.36) | 2.221 | 2.082 | (–6.67) |
| 4 | 1.739 | 1.710 | (–1.72) | 2.686 | 2.411 | (–11.39) |
| 8 | 1.829 | 1.803 | (–1.43) | 3.047 | 2.947 | (–3.42) |
| Initial in-sample estimation: 1981:1–89:4 | | | | | | |
| Forecasting periods: 1990:1–94:4 | | | | | | |
| (CPIC) | | | | | | |
| 1 | .879 | .814 | (–8.07) | 2.185 | 2.086 | (–4.77) |
| 4 | .866 | .785 | (–10.33) | 1.901 | 2.168 | (12.35) |
| 8 | .856 | .710 | (–20.65) | 1.893 | 2.260 | (16.24) |
| (CPI) | | | | | | |
| 1 | 1.339 | 1.268 | (–5.58) | 2.213 | 2.086 | (–.88) |
| 4 | 1.568 | 1.439 | (–8.95) | 2.188 | 2.168 | (–.88) |
| 8 | 1.328 | 1.172 | (–13.26) | 2.099 | 2.260 | (7.12) |
| Initial in-sample estimation: 1957:1–77:4 | | | | | | |
| Forecasting periods: 1978:1–81:4 | | | | | | |
| (CPIC) | | | | | | |
| 1 | 2.836 | 3.085 | (8.08) | 3.987 | 3.859 | (–3.33) |
| 4 | 3.102 | 3.564 | (12.97) | 5.017 | 4.442 | (–12.93) |
| 8 | 4.041 | 4.206 | (3.92) | 4.313 | 4.838 | (10.86) |
| (CPI) | | | | | | |
| 1 | 2.413 | 2.358 | (–2.31) | 2.930 | 3.859 | (24.07) |
| 4 | 4.102 | 3.967 | (–3.41) | 5.526 | 4.442 | (–24.39) |
| 8 | 4.925 | 5.038 | (2.24) | 4.823 | 4.838 | (.33) |

NOTES: The forecasts for the ECM were formed using VAR(4) (including error-correction term and a constant). The entries in the table refer to the root mean square forecast error. To forecast inflation and wage growth for out-of-sample periods, we reestimate both ECM and AR model by updating the in-sample periods. For example, the four-quarter-ahead forecast for inflation and wage growth for 1991:1 is constructed by the models estimated over the period 1957:1–90:1 or 1981:1–90:1.

Figure 3
**Four-Quarter Ahead Out-of-Sample Forecast Of CPI Core Inflation, 1990–94**

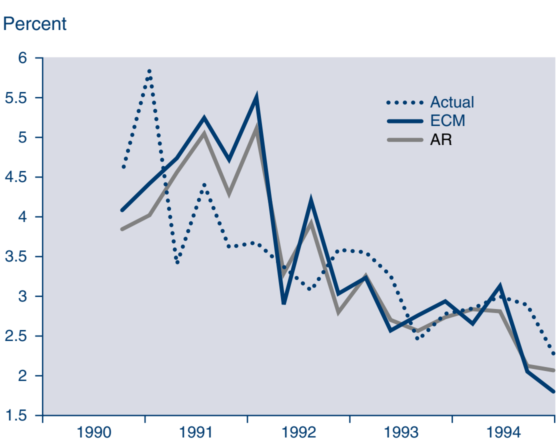
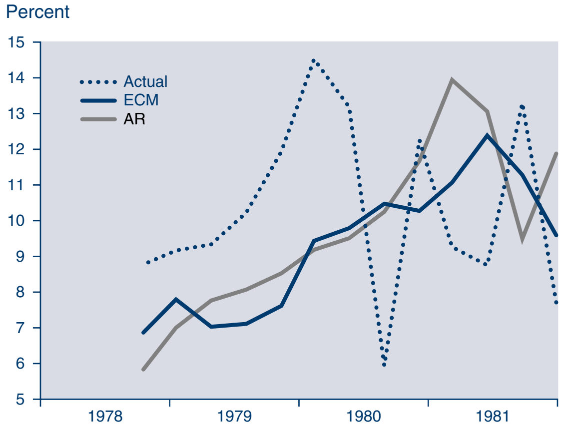


Figure 4

**Four-Quarter Ahead Out-of-Sample Forecast Of CPI Core Inflation, 1978–81**




In summary, out-of-sample forecasts offer little support that the growth of unit labor costs substantially helps forecast inflation, especially in recent years. For forecasting inflation during the 1990s, of the models we consider, a univariate autoregressive model of inflation using only post-1980 data results in the smallest forecast errors. The out-of-sample forecasts for the late 1970s indicate that wage growth did modestly help to forecast CPIC inflation.

## Conclusions

Many analysts have heralded the slow growth of unit labor costs during recent years as a harbinger of continued low inflation. In this article, we investigate the usefulness of labor costs as a predictor of inflation. Earlier studies have focused on in-sample causality tests. Our in-sample causality tests indicate that, during the pre-1980 period, wage growth did have information content for future core inflation (CPIC) but not overall CPI inflation. During the post-1980 period, however, this information content has disappeared. Additionally, we find that the evidence of inflation causing wage growth is quite robust across samples.

In contrast with earlier studies, we also investigate out-of-sample forecasts of inflation using labor costs in an error-correction model. Out-of-sample forecasts offer the ultimate test of whether wages help predict future inflation. For recent years, the out-of-sample forecasting exercises offer no evidence that wage growth contributes to any reduction in forecast errors compared with univariate autoregressive models of inflation. Therefore, when assessing future inflation developments, these results suggest that policymakers and analysts should put little weight on recent wage trends.

## Notes

We would like to thank Nathan Balke, Joseph Haslag, and Evan Koenig for helpful comments and suggestions. Any remaining errors are our own.

1 The Granger causality test is simply a statistical methodology for showing whether a variable contains information about subsequent movements in another variable.

2 However, Mehra finds that the presence of this bidirectional causality is sensitive to how inflation is modeled.

3 The motivation for Mehra's work is to examine the hypothesis that prices are marked up over productivity-adjusted labor costs, a central proposition of the expectations-augmented Phillips curve model. If that hypothesis is correct, then long-run movements in prices and labor costs must be correlated, and short-run movements in labor costs should help predict short-run movements in prices. Therefore, Mehra's results are consistent with the markup hypothesis for consumer prices but not for the implicit price deflator.

4 Any instability in the wage-price relationship could also be a source of instability in the price markup hypothesis and the Phillips curve (see Mehra 1993).

5 Unit labor costs are for the nonfarm business sector. The figure for consumer prices including food and energy is qualitatively similar. The formal analysis in this study is carried out using both measures of consumer prices.

6 The 1980:4 breakpoint was chosen arbitrarily on the basis of looking at Figure 1.

7 We use consumer prices both because they are perhaps the most closely watched measure of underlying inflation and because of Mehra's finding that labor costs do have information content for future consumer prices.

8 A time series is nonstationary if it has a time-varying mean and/or variance. Nonstationarity of a series violates an assumption underlying many statistical inferences and can lead to "spurious regression phenomenon," first described by Granger and Newbold (1974). One commonly used way of removing nonstationarity is to take first differences of the series.

9 In other words, second differencing is required for stationarity. Mehra (1991) reports similar results, while Mehra (1993) finds consumer prices to be I(1). Throughout the analysis that follows, we check the sensitivity of our results to the finding that prices are I(2).

10 The concept of cointegration, first proposed by Granger and Weiss (1983), is fundamental to the use of the error-correction model. Engel and Granger (1987) show that a model estimated using differenced data will be misspecified if the variables are cointegrated and the cointegrating relationship is ignored. Cointegration of two series means that they are nonstationary and tend to move together such that a linear combination of them is stationary. Cointegration is sometimes interpreted as representing a long-run equilibrium (steady-state) relationship.

[11] Since the standard errors are corrected for serial correlation, the $\chi^2$ statistic is appropriate to test whether the $\alpha$'s and $\beta$'s are significant.

[12] The ADF test is applied to the long-run relationship. In other words, $\Delta p - \alpha_p - \beta_p \Delta w$ and $\Delta w - \alpha_w - \beta_w \Delta p$.

[13] The Schwartz information criterion always implies a lag length of four or less. Since the results are not sensitive to the choice of lag length ($k$ = 2, 4, or 8), we report only those from the model of $k$ = 4.

[14] This result is consistent with Mehra (1991), which models consumer prices as I(2). However, Mehra (1993) models prices as I(1) and finds significant causality. As a robustness check, we also find causality for the whole sample and both price measures if we model prices as I(1).

[15] This maximum $\chi^2$ statistic is sometimes called the Quandt likelihood-ratio statistic, which tests for a break in any or all of the coefficients.

[16] Formally, the rejection of the null of a stable cointegrating vector implies two alternatives: no cointegration and therefore no error-correction model, or an error-correction model in which there is assumed to be two cointegrating vectors, one from each subsample.

[17] Another important implication of Hansen's test is that the lack of cointegration is a special case of the alternative hypothesis, so the SupF test can also be viewed as a test of the null of cointegration against the alternative of no cointegration. If the SupF rejects the null, one may conclude that the standard model of cointegration, including its implicit assumption of long-run stability of the cointegrating relationship, is rejected by the data.

[18] For both CPIC and CPI, the results for the subsamples do not change when prices are modeled as I(1). Thus, only the whole sample results for the CPI are sensitive to whether prices are modeled as I(1) or I(2).

[19] Formally, the results from the two subsamples are conditioned on there being a structural break in the ECM. However, it should be noted that rolling formal stability tests cannot reject the null hypothesis of stability. Our choice to examine the results for the post-1980 sample was made on the basis of rejecting stability of the cointegrating relationship. Additionally, the results from the Baxter and King filter analysis and Figure 1 motivated us to examine the results for the post-1980 period.

[20] The ECM and autoregressive models with four lags perform superior to alternative lag lengths.

[21] Encompassing tests due to Chong and Hendry (1986) confirm this finding.

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