WALTER WASSERFALLEN

# Trends, Random Walks, and the Expectations-Augmented Phillips Curve Evidence from Six Countries

IN MOST MACROECONOMIC MODELS, variations in nominal variables, such as inflation or money growth, are considered to be important determinants of cyclical fluctuations in real activity. The most prominent hypothesis in this respect is the so-called Phillips curve. In its modern interpretation, this hypothesis maintains that unexpected changes in nominal magnitudes lead to variations in real variables. Empirical work in this area, however, is quite difficult, in part because both cyclical output movements and unexpected changes in nominal magnitudes are unobservable. Such series must therefore be approximated through procedures that are ad hoc in many important respects.

The empirical implementation of business cycle models, such as the expectations-augmented Phillips curve, requires assumptions with respect to the decomposition of time series into growth and seasonal components in order to isolate cyclical movements and to avoid misspecified equations. In this endeavor, it has become general practice to assume that economic growth can be reasonably well approximated by a deterministic linear time trend. Seasonality may then be cap-

The comments by Ernst Baltensperger, Charles Bean, Stephen Cecchetti, Michael Driscoll, Martin Eichenbaum, David Laidler, Bennett McCallum, Allan Meltzer, Frederic Mishkin, Michael Parkin, René Stulz, and participants at seminars held at Carnegie-Mellon University, Columbia University, Ohio State University, Universität Bern, University of Western Ontario, Universität Basel and at the Seminaire d'économie monétaire internationale Paris are gratefully acknowledged. The help received from two referees and the editors has been especially helpful in preparing the final version.

WALTER WASSERFALLEN *is lecturer and deputy director, Study Centre Gerzensee, Foundation of Swiss National Bank.*

*Journal of Money, Credit, and Banking*, Vol. 20, No. 3 (August 1988, Part 1)

tured through the explicit introduction of dummy variables or through the use of seasonally adjusted data.[1] Recent investigations into the stochastic characteristics of macroeconomic time series by Nelson and Plosser (1982), Stulz and Wasserfallen (1985), and Wasserfallen (1986) as well as results presented in the next section provide evidence that variables measuring real activity contain unit roots. This finding suggests that the growth and seasonal parts of real activity variables may be more adequately modeled by appropriate random walks.

If this conclusion is valid, regressions of real activity measures on a constant, a linear time trend, seasonal dummies, and proxies for expected and unexpected changes in nominal magnitudes most likely yield misleading results. Chan, Hayya, and Ord (1977), Nelson and Kang (1981, 1984), and Phillips (1986) show that the regression of a random walk variable on a linear trend produces nonstationary residuals, which is inconsistent with the assumptions necessary for standard statistical inference procedures. In finite samples, the estimated residuals are positively but spuriously autocorrelated, generating the wrong impression of a highly persistent cyclical component. The explanatory power of time and other exogenous variables can furthermore be vastly overstated. Studies by Darby (1983), Haraf (1983), Makin (1982), and Wasserfallen (1985) actually point in the direction that the positive evidence for the Phillips curve often reported in the literature may be an artificial result due to inappropriate assumptions concerning growth and seasonal components.

This impression is confirmed here using quarterly observations from six countries—the United States, France, West Germany, Great Britain, Italy, and Switzerland—over the post–World War II period. The influence of unanticipated changes in nominal magnitudes is quite strong when growth and seasonality are respectively measured by a linear trend and seasonal dummies. On the other hand, virtually no effects are found if the regressions are estimated in the appropriate differences of real activity variables.

The paper is organized as follows: Test results for nonstationarity in the mean of variables measuring real activity are presented in the next section. The models underlying the Phillips curve estimates and the econometric techniques are discussed in section 2. Section 3 contains the empirical findings. A summary and some conclusions complete the paper.

## 1. NONSTATIONARITIES IN REAL ACTIVITY

It is common knowledge that variables such as real GNP or industrial production have a tendency to grow over time, notably in industrial countries over the post–World War II period. In the following, two widely used possibilities to

[1]Examples of this methodology from the Phillips curve literature include Attfield, Demery and Duck (1981), Barro (1978), Barro and Rush (1980), Blejer and Leiderman (1980), Cuddington (1980), Evans (1984), Fair (1979), Gordon (1982), Hoffman and Schlagenhauf (1982), Lawrence (1983), Leiderman (1979), Lucas (1973), Mishkin (1982), Sargent (1976), Sheffrin (1979), and Taylor (1979).

account for this nonstationarity in the mean are empirically evaluated. The first alternative is a deterministic linear time trend, generally introduced in empirical work on the expectations-augmented Phillips curve. The second hypothesis, popularized by Box and Jenkins (1976), maintains that the series contain unit roots and that appropriate differencing is therefore the correct way to remove nonstationarities. With respect to seasonality, the analogous distinction is between seasonal dummy variables and seasonal differencing.

Test procedures recently developed by Fuller (1976) and Hasza and Fuller (1982) allow a discrimination between these competing hypotheses. The basic idea is to isolate possible nonstationarities through coefficients in a regression equation, assumed to be of finite autoregressive form. Conventional $t$- and $F$-statistics are then calculated to test the hypotheses of interest. The usual significance levels, however, are inappropriate if unit roots are present and must be replaced through the relevant test statistics contained in the papers mentioned above. The general model used in the empirical work is

$$y_t = \beta_0 + \beta_1 t + \sum_{j=1}^{s-1} \gamma_j D_{jt} + \alpha_1 y_{t-1} + \alpha_2 (y_{t-s} - y_{t-s-1})$$
$$+ \alpha_3 (y_{t-1} - y_{t-s-1}) + k_1 w_{t-1} + \ldots + k_2 w_{t-8} + \epsilon_t \; . \qquad (1)$$

$y$ is the real activity variable under investigation and $t$ a deterministic time trend. The $D_j$s are seasonal dummies with $s$ the seasonal frequency, which will be 4 in the applied work executed with quarterly data. The $w$s, defined as $w_t = (1-L)(1-L^s)y_t$, are included to capture remaining autocorrelation patterns so that $\epsilon$ is a well-behaved error term. The lag operator $L$ is given by $L^j y_t = y_{t-j}$.

The performed tests are summarized in Table 1. Procedure I is concerned only with the presence of a first-order unit root assuming the seasonal part to be stationary. Model II tests for first- as well as seasonal-order unit roots. The appropriate restrictions to evaluate the existence of deterministic trend and seasonal elements are formulated in hypothesis III.

The results obtained for seasonally unadjusted quarterly time series of indus-

TABLE 1

TESTS FOR NONSTATIONARITY

| Test | Restrictions tested in equation (1) | Test statistic | Significance table |
|---|---|---|---|
| I | $\alpha_1 = 1$ | $\hat{\tau}_\tau$ | Fuller (1976, Table 8.5.2) |
| II | $\alpha_1 = \alpha_2 = 1, \alpha_3 = 0$ | $\theta^3_{n-d-4}$ | Hasza and Fuller (1982, Table 5.1) |
| III | $\alpha_1 = \alpha_2 = 1$, $\alpha_3 = \beta_0 = \beta_1 = \gamma_j$ (all $j$) $= 0$ | $\theta^{d+4}_{n-d-4}$ | Hasza and Fuller (1982, Table 5.1) |

trial production are displayed in Table 2.[2] Natural logarithms of the original data are used throughout. The row labeled I clearly reveals that the null hypotheses $\alpha_1 = 1.0$ cannot be rejected for all countries and that the parameter estimates are generally very close to one. The outcome for hypothesis II is equally favorable. The assumption that first as well as seasonal-order unit roots are present is only violated for the United States and Great Britain. Hypothesis III, testing for the simultaneous occurrence of deterministic and stochastic trends, is rejected for the United States only on the 5 percent level but not at all on the 10 percent level. Given these results, the conclusion must be that nonstationarities in real activity measures are more adequately modeled by unit roots than by deterministic components.

TABLE 2

NONSTATIONARITIES IN REAL ACTIVITY

| | United States | France | West Germany | Great Britain | Italy | Switzerland |
|---|---|---|---|---|---|---|
| Observation period | 1965.II–1984.I | 1973.II–1983.II | 1966.II–1984.I | 1963.II–1983.I | 1976.IV–1984.I | 1965.II–1984.I |
| Unit root tests | | | | | | |
| I | 0.75 (−2.65) | 0.43 (−2.20) | 0.95 (−1.02) | 0.92 (−1.22) | 0.69 (−0.68) | 0.91 (−1.52) |
| II | 11.46* | 6.60 | 6.37 | 9.81* | 3.44 | 5.92 |
| III | 4.55* | 3.18 | 2.59 | 4.31 | 1.81 | 2.40 |
| Deterministic Trends | | | | | | |
| $t$ | 0.007* (24.49) | 0.0008* (1.80) | 0.006* (13.73) | 0.003* (11.47) | 0.001 (1.19) | 0.004* (10.12) |
| Seasonal dummies | 0.16 | 99.59* | 3.16* | 11.03* | 24.91* | 5.63* |
| $r_1$ | 0.88* | 0.78* | 0.93* | 0.87* | 0.87* | 0.88* |
| $r_2$ | 0.67* | 0.54* | 0.83* | 0.76* | 0.69* | 0.82* |
| $r_3$ | 0.45* | 0.23 | 0.73* | 0.63* | 0.46 | 0.70* |
| $r_4$ | 0.24 | −0.03 | 0.61* | 0.51* | 0.24 | 0.66* |
| Differences | | | | | | |
| $r_1$ | 0.50* | 0.29 | 0.24 | −0.05 | 0.20 | −0.05 |
| $r_2$ | 0.12 | 0.25 | 0.11 | 0.04 | 0.20 | 0.08 |
| $r_3$ | −0.13 | −0.20 | 0.09 | 0.28* | −0.16 | −0.12 |
| $r_4$ | −0.41* | −0.53* | −0.33* | −0.51* | −0.33 | −0.38* |

NOTES: The dependent variable is the natural logarithm of industrial production. Observation period is in quarters.
Unit root tests: Tests as defined in Table 1 based on ordinary-least-squares estimates of equation (1). The $t$-value in test I is relative to $\alpha_1 = 1.0$.
Deterministic trends: Ordinary-least-squares estimates of dependent variable in linear time trend $t$ and seasonal dummies ($F$-value for joint significance is shown)
Differences: First and seasonal differences of dependent variable
$r_j$: Estimated autocorrelation at lag $j$
$t$-values are shown in parentheses below the estimated coefficients.
* indicates the estimated coefficient is significant on the 5 percent level.

[2]As an alternative, real GNP has also been analyzed for several countries with virtually identical findings. The same is true with respect to the Phillips curve estimates presented in section 3.

It must be mentioned at this point that the test procedures used above have virtually no power to discriminate between a unit root and a corresponding autocorrelation coefficient slightly lower than one combined with a deterministic trend component.[3] Obviously, the implications for business cycle measurement are crucial. If a unit root is present, the series is nonstationary, whereas in the other case it would finally return to the trend line and would therefore exhibit cyclical swings of long duration. Stulz and Wasserfallen (1985, p. 14), however, note that three characteristics of the empirical results favor the unit root interpretation. First, the estimates available in the literature indicate that the respective parameter values are independent of the measurement interval. Given the length of a possible cyclical adjustment period, they should, however, be inversely related to the time interval between observations. Second, the implied length of the business cycle appears to be implausibly long on economic grounds if the $\alpha_1$-parameter is actually around 0.9. Third, the uniformity of the estimates both across countries and time series measuring real activity is also in favor of the unit root interpretation, unless the speed of adjustment is equally low in all markets and countries.

Table 2 provides two additional pieces of information relevant to the purposes of this paper. First, the results of regressing the level of industrial production on a constant, a linear time trend, and seasonal dummies are shown under the heading "Deterministic Trends." The trend and the dummies are highly significant. $R^2$, measuring the explanatory power, is around 0.9 in all cases. Highly serially correlated residuals are another interesting regularity indicating strong persistence in the cyclical component. Nelson and Kang (1984) demonstrate that these are exactly the characteristics one would expect if the time series under investigation contain unit roots. They also show that the residuals are in fact nonstationary and that the measured autocorrelation is a pure statistical artifact, depending exclusively on sample size. The autocorrelation structure of the first and seasonally differenced series, shown in Table 2 under "Differences," leads to almost opposite conclusions. Only a few coefficients are larger than two standard errors. A notable feature is the sometimes large negative value at the first seasonal lag, indicating seasonal overdifferencing. Interestingly, the serial correlation coefficients at seasonal lags damp out only very slowly in the seasonally undifferenced series. These findings imply that the persistence of the stationary cyclical component might be quite small. Given this empirical outcome, it must be asked whether estimates of the expectations-augmented Phillips curve are sensitive with respect to the assumptions made about growth and seasonal components in real activity variables. The analysis and results presented in the next sections will, among other things, provide an answer to this questions.

[3]McCallum (1986) provides a detailed discussion of these issues in the context of business cycle models.

## 2. MODELS AND ECONOMETRIC PROCEDURES

In this section, the theoretical foundation for the subsequent empirical investigations is laid. Starting point is the additive decomposition of the real activity variable $y$ in its trend or growth, cyclical, and seasonal components, denoted respectively by $T$, $C$, and $S$.

$$y_t = T_t + C_t + S_t \tag{2}$$

where $t$ is a time index, measured in quarters. Following the general understanding in macroeconomic theory, the cyclical part is assumed to possess a constant long-run mean of zero. Nonstationarities in real activity variables are therefore attributed to the trend and seasonal components, respectively.

The two approaches with respect to $T$ and $S$ mentioned in the previous sections are formalized in equations (3)–(6). In most of the existing empirical investigations, $T$ is assumed to follow a deterministic linear trend and the seasonal part is modeled through appropriate dummy variables.[4] The resulting equations are

$$T_t = \beta_0 + \beta_{1t} ; \tag{3}$$

$$S_t = \gamma_1 D_{1t} + ..... + \gamma_{s-1} D_{s-1t} . \tag{4}$$

$s$ is the seasonal periodicity of the data, which is 4 in the empirical work presented below.[5]

The alternative formulation, based on stochastic trends for the growth and seasonal components, is given by equations (5) and (6). $T$ is assumed to follow a random walk with drift, which allows the maximum possible degree of persistence for the cyclical part. The seasonal process also contains a unit root plus a stationary part, capturing any remaining autocorrelation at the seasonal lag.[6]

$$T_t = T_{t-1} + K + \eta_{1t} ; \tag{5}$$

$$S_t = S_{t-s} + \eta_{2t} - \Delta_1 \eta_{2t-s} . \tag{6}$$

The model for the cyclical component is common to both approaches. It states that $C_t$ is dependent on a set of exogenous magnitudes plus a possibly autocorrelated random influence, yielding

[4]Alternatively, seasonally adjusted data are used. This approach is not pursued in this paper. Note further that a stochastic term might be added to equations (3) and (4) without changing the relationships derived for estimation purposes.

[5]$s-1$ seasonal dummies together with the constant term are introduced in order to avoid perfect multicollinearity in estimation.

[6]In practice, only $\eta_{2t-s}$ plays a role but the generalization to higher-order processes would be straightforward.

$$\begin{aligned} C_t &= f_0 x_t + f_1 x_{t-1} + f_2 x_{t-2} + .. + \eta_{3t} - \Theta_1 \eta_{3t-1} - .. \\ &= f(L)x_t + \Theta(L)\eta_{3t} \ . \end{aligned} \tag{7}$$

The $x$ variables are of course given by the hypothesis under investigation. In the context of this study, $x$ denotes unexpected changes in nominal variables, specifically inflation or monetary growth.

The test formulations are now easily derived. The traditional approach is given by the combination of equations (2), (3), (4), and (7), yielding

$$y_t = \beta_0 + \beta_1 t + \sum_{j=1}^{s-1} \gamma_j D_{jt} + f(L)x_t + \Theta(L)\eta_{3t} \ . \tag{8}$$

The alternative formulation follows from equations (2), (5), (6), and (7). The result, stated in the stationary differences of the endogenous variable, is

$$\begin{aligned} Dy_t &= (1-L^s)\eta_{1t} + (1-L)(\eta_{2t} - \Delta_1 \eta_{2t-1}) + f(L)Dx_t + D\eta_{3t} \\ &= g(L)x_t + \delta(L)\epsilon_t \ . \end{aligned} \tag{9}$$

$D$ is equal to $(1-L)(1-L^s)$ and $g(L) = f(L)D$. Equations (8) and (9) form the basis of the empirical work reported in the following section. For estimation purposes, the terms containing the unobservable random variables $\eta_{1t}$, $\eta_{2t}$, and $\eta_{3t}$ are combined into a single error process, $\delta(L)\epsilon_t$. Note further, that the drift term in the growth component, $K$, vanishes due to seasonal differencing of $y_t$. Equations (8) and (9) are similar to the specifications used in many other studies. Specifically, unexpected changes in nominal magnitudes are assumed to be exogenous relative to other determinants of real activity and to real activity itself. Under these circumstances, the error term becomes independent of the $x$s and the use of single equation estimation techniques can consequently be justified.[7]

In the empirical work, nominal magnitudes are alternatively measured by the rate of inflation, as implied by consumer price indices, and the growth rate of the money stock M1. Unanticipated variations in nominal variables, the $x$s in equations (8) and (9), are proxied by the estimated in-sample residuals from appropriate ARIMA models.[8]

The two-step procedure developed by Barro is used to estimate the various Phillips curves. First, the residuals from the forecasting schemes with respect to

[7]Admittedly, these are strong assumptions but there is no obvious reason why they should bias the comparison between specifications in levels and differences.

[8]Various other specifications have been tried as well. Inflation surprises have been measured through the residuals of transfer functions relating inflation to the growth rate of M1 and the monetary base or through innovations in real Euromarket interest rates. The monetary base has been considered as an alternative to M1. Fixed and flexible exchange rate periods have furthermore been distinguished. The conclusions derived from the empirical results presented in the next section are not affected by these choices. The respective estimates are therefore not reported.

nominal magnitudes are calculated. In the second step, these estimates are introduced as exogenous variables in the Phillips curve regressions. The transfer function methodology developed by Box and Jenkins (1976) is applied in order to account for serial correlation in the residuals. In practice, an ARMA model for the error term is specified and estimated based on the observed autocorrelation structure of the estimated residuals from a usual OLS regression.

Hoffman, Low, and Schlagenhauf (1984) and Pagan (1984) show that this procedure generates consistent parameter estimates but that test statistics like $t$- and $F$-values are overstated because generated regressors are used. The null hypothesis of no influence of unanticipated changes in nominal variables on real activity will be rejected too often under these circumstances. Compared to FIML estimators proposed by Abel and Mishkin (1983), the advantage of the two-step approach is its simplicity. Hoffman et al. (1984) derive a corrected test statistic based on the two-step procedure which is, however, quite complicated as well. For two reasons, the conventioal two-step estimator will nevertheless be used in this paper. First, the empirical results provide evidence in favor of no significant Phillips curve relationships. This outcome would obviously be strengthened if the biases discussed above would be taken into account.[9] Second, the empirical comparison of different Phillips curve specifications should not be affected by using Barro's original procedure.

## 3. EMPIRICAL FINDINGS

Empirical estimates for the real effects of unexpected inflation and unanticipated variations in M1 are presented in Tables 3 and 4 respectively. The contemporaneous and four-lagged terms of the exogenous variable are included in the transfer functions.

In Table 3, the real effects of unexpected inflation are compared for equations (8) and (9) in order to show clearly the differences in results that might arise between a deterministic and a stochastic trend assumption. For four out of six countries, the Phillips curve is apparently well confirmed by the regressions including deterministic trends. Great Britain and Switzerland are exceptions but the distributed lag coefficients on unexpected inflation are mostly positive in these cases also. The time trend and the seasonal dummies (not shown) generally appear to be important determinants. Furthermore, the explanatory power is quite high. For all countries, residual autocorrelation can be removed through a first-order autoregressive process. The respective coefficient is very close to one throughout, indicating that unexpected inflation cannot account for the high serial correlation in the business cycle proxy obtained by assuming deterministic growth and seasonal components in real activity.

[9]Note in this context that Hoffman and Schlagenhauf (1982) obtain significant influences of unexpected changes in the money supply on real activity for several countries by using the FIML procedure combined with a deterministic time trend proxying for real growth.

TABLE 3

REAL EFFECTS OF UNANTICIPATED INFLATION

| | United States Trend | United States Diff. | France Trend | France Diff. | West Germany Trend | West Germany Diff. | Great Britain Trend | Great Britain Diff. | Italy Trend | Italy Diff. | Switzerland Trend | Switzerland Diff. |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Observation Period | 1964.III–1984.I | | 1970.II–1983.II | | 1969.II–1984.I | | 1960.I–1983.I | | 1974.IV–1983.IV | | 1963.II–1984.I | |
| Exogenous Variables | | | | | | | | | | | | |
| Constant | 4.23* | 0.00 | 4.59 | −0.01 | 5.19* | 0.00 | 4.24* | 0.00 | 4.65* | 0.01 | 5.02* | 0.00 |
| | (75.62) | (1.29) | (93.05) | (−0.39 | (81.18) | (−1.41) | (48.70) | (0.21) | (43.50) | (0.12) | (64.41) | (0.00) |
| $t$ | 0.008* | | 0.003* | | 0.004* | | 0.004* | | 0.005* | | 0.005* | |
| | (10.62) | | (3.50) | | (4.88) | | (4.05) | | (1.91) | | (4.12) | |
| $P^u_t$ | 3.02* | 1.33* | 1.30* | −0.09 | −0.18 | −1.01 | −0.24 | 0.10 | 0.83* | −0.24 | −0.44 | −0.74 |
| | (5.66) | (2.13) | (2.83) | (−0.16) | (−0.23) | (−1.23) | (−1.04) | (0.43) | (1.97) | (−0.42) | (−0.74) | (−1.27) |
| $P^u_{t-1}$ | 2.80* | −0.79* | 1.83* | −0.04 | 1.49 | −1.64* | 0.04 | −0.21 | 1.55* | −1.06* | 0.52 | −1.10 |
| | (4.44) | (−1.91) | (3.02) | (−0.08) | (1.45) | (−2.33) | (0.12) | (−1.50) | (3.19) | (−1.86) | (0.68) | (−2.04) |
| $P^u_{t-2}$ | 2.35* | −1.71* | 1.55* | −0.52 | 2.41* | 0.14 | 0.23 | −0.23 | 0.53 | −2.14* | 0.92 | −0.22 |
| | (3.52) | (−4.37) | (2.42) | (−0.96) | (2.12) | (0.21) | (0.69) | (−1.59) | (1.09) | (−3.73) | (1.07) | (−0.49) |
| $P^u_{t-3}$ | 1.66* | −0.97* | 1.13* | −0.21 | 2.88* | −1.13 | 0.17 | 0.10 | −0.34 | −1.25* | 1.16 | −0.01 |
| | (2.63) | (−2.37) | (1.81) | (−0.40) | (2.74) | (−1.59) | (0.57) | (0.69) | (−0.72) | (−2.22) | (1.45) | (−0.02) |
| $P^u_{t-4}$ | 0.69 | −3.13* | 0.11 | −1.25* | 1.49* | 0.03 | 0.16 | −0.17 | −0.51 | −0.41 | 0.15 | 0.27 |
| | (1.30) | (−4.96) | (0.22) | (−2.22) | (1.81) | (0.03) | (0.73) | (−0.72) | (−1.40) | (−0.73 | (0.23) | (0.47) |
| Error process | | | | | | | | | | | | |
| $\phi_1$ | 0.89* | 0.21 | 0.79* | | 0.85* | | 0.94* | | 0.93* | | 0.89* | |
| | (17.07) | (1.57) | (8.55) | | (11.74) | | (27.12) | | (17.02) | | (18.63) | |
| $\Delta_1$ | | 1.01* | | 0.58* | | 0.78* | | 0.82* | | | | 0.71* |
| | | (14.17) | | (4.22) | | (8.09) | | (10.09) | | | | (7.06) |
| Test statistics | | | | | | | | | | | | |
| $\bar{R}^2$ | 0.99 | 0.66 | 0.99 | 0.31 | 0.99 | 0.35 | 0.97 | 0.37 | 0.99 | 0.38 | 0.96 | 0.31 |
| $r_1$ | 0.24 | −0.03 | 0.05 | 0.12 | 0.00 | −0.10 | −0.01 | −0.02 | 0.21 | 0.20 | −0.20 | −0.17 |
| $r_2$ | 0.07 | 0.09 | −0.11 | −0.04 | −0.01 | −0.21 | 0.12 | 0.05 | −0.09 | 0.16 | 0.21 | 0.08 |
| $r_3$ | −0.03 | −0.09 | 0.17 | −0.09 | 0.24 | 0.07 | 0.03 | 0.10 | 0.06 | −0.08 | −0.16 | −0.15 |
| $r_4$ | −0.08 | 0.02 | 0.17 | −0.05 | −0.08 | −0.04 | 0.22 | 0.02 | −0.12 | −0.26 | 0.23 | 0.01 |

NOTES: Transfer function estimates of equations (8) and (9).
Trend: The dependent variable is the natural logarithm of industrial production. Seasonal dummies are included but not shown
Diff.: The dependent variable is the natural logarithm of industrial production first and seasonally differenced
$t$: Linear time trend
$P^u$: Unexpected inflation, measured as residuals of ARIMA model for consumer price index
Error process: ARMA model for the error term with $\phi_1$ = first-order autoregressive coefficient and $\Delta_1$ = first-order seasonal moving average coefficient
$\bar{R}^2$: $R^2$, adjusted for degrees of freedom
$r_j$: Autocorrelation coefficient of residuals at lag $j$
$t$-values are shown in parentheses below the estimated coefficients. * indicates that coefficients are significantly different from zero on the 5 percent level.

TABLE 4

REAL EFFECTS OF UNEXPECTED CHANGES IN THE MONEY SUPPLY

| | United States | France | West Germany | Great Britain | Italy | Switzerland |
|---|---|---|---|---|---|---|
| Observation period | 1964.III–1984.I | 1974.II–1983.II | 1964.II–1984.I | 1965.II–1983.I | 1974.IV–1983.IV | 1964.I–1984.I |
| Exogenous variables | | | | | | |
| Constant | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | (−1.21) | (−0.96) | (−0.68) | (−1.19) | (−0.57) | (−0.41) |
| $M^u_t$ | 0.00 | −0.54 | 0.00 | −0.09 | 0.80* | 0.11 |
| | (−0.02) | (−1.62) | (0.00) | (−0.72) | (2.73) | (0.57) |
| $M^u_{t-1}$ | 0.39* | 0.30 | 0.02 | 0.03 | 0.21 | −0.26* |
| | (2.15) | (1.15) | (0.14) | (0.28) | (0.93) | (−2.10) |
| $M^u_{t-2}$ | −0.08 | −0.08 | 0.34* | 0.08 | −0.13 | 0.05 |
| | (−0.39) | (−0.33) | (2.14) | (0.84) | (−0.59) | (0.40) |
| $M^u_{t-3}$ | 0.41* | 0.25 | 0.24 | −0.01 | 0.19 | −0.03 |
| | (2.16) | (0.94) | (1.49) | (−0.07) | (0.83) | (−0.24) |
| $M^u_{t-4}$ | −0.17 | 0.59 | −0.07 | 0.20 | −0.12 | 0.16 |
| | (−0.64) | (1.71) | (−0.31) | (1.51) | (−0.40) | (0.77) |
| Error process | | | | | | |
| $\phi_1$ | 0.42* | | | | | |
| | (3.81) | | | | | |
| $\Gamma_1$ | | −0.56* | | | | |
| | | (−4.09) | | | | |
| $\Delta_1$ | 1.04* | | 0.67* | 0.77* | 0.75* | 0.92* |
| | (18.53) | | (6.94) | (8.04) | (5.10) | (14.91) |
| Test statistics | | | | | | |
| $\bar{R}^2$ | 0.60 | 0.27 | 0.28 | 0.36 | 0.40 | 0.30 |
| $r_1$ | 0.09 | 0.17 | 0.11 | −0.02 | 0.12 | −0.17 |
| $r_2$ | −0.15 | 0.13 | 0.12 | 0.09 | 0.03 | 0.08 |
| $r_3$ | 0.04 | −0.06 | 0.10 | 0.10 | −0.28 | −0.13 |
| $r_4$ | 0.03 | −0.19 | 0.05 | −0.02 | −0.05 | 0.17 |

NOTES: Transfer function estimates of equation (9). The dependent variable is the natural logarithm of industrial production first and seasonally differenced.
$M^u$: Unexpected growth rate of M1, measured as residuals of ARIMA model
Error process: ARMA model for the error term with $\phi_1$ = first-order autoregressive coefficient, $\Gamma_1$ = first-order seasonal autoregressive coefficient, and $\Delta_1$ = first-order seasonal moving average coefficient.
$r_j$: Autocorrelation coefficient of residuals at lag $j$
$t$-values are shown in parentheses below the estimated coefficients. * indicates that coefficients are significantly different from zero on the 5 percent level.

Given the stochastic characteristics of industrial production discussed in section 1, first and seasonal differencing of the endogenous variable appears to be the more appropriate procedure to proxy for growth and seasonality and to achieve stationarity. The estimates incorporating this hypothesis exhibit distinctively different characteristics. The influence of unexpected inflation on real activity vanishes almost completely or becomes even negative, contrary to the usual Phillips curve hypothesis. Considerable changes in the error process are also observed, with a seasonal moving average parameter now becoming the dominant element.[10] Note further that the explanatory power is substantially reduced.

Almost the same results are found for the effects of unanticipated variations in M1 on industrial production using the specification in differences. Table 4 reveals that only isolated coefficients for the United States, West Germany, and Italy support the hypothesized positive influence. The structure of the error terms is also strikingly similar.[11]

## CONCLUSIONS

In the preceding sections, two major empirical results are reported. First, the growth and seasonal components of real activity observed in six industrialized countries over the post–World War II period can be characterized more adequately by appropriate random walks than by deterministic trends. Second, the influence of unexpected changes in the rate of inflation and the growth rate of the money stock M1 on real activity is evaluated. It turns out that the distinction between deterministic and stochastic trends, used to decompose real activity into growth, cyclical, and seasonal parts, affects the outcome of these tests quite strongly. The Phillips curve can be confirmed assuming deterministic trends. The most likely reason for this outcome is that this procedure leads to a highly but spuriously autocorrelated proxy for the business cycle making it relatively easy to obtain significantly positive parameters in a distributed lag regression. The indicated persistence in the cyclical part is greatly reduced if differencing is applied to achieve stationarity. In most cases, the real effects of unforeseen movements in nominal variables are reduced to zero or become even negative under these circumstances. This is additional evidence that the spurious regression phenomenon discussed by Granger and Newbold (1974) is an important issue in empirical work.[12]

*Data for this paper are available at the* JMCB *editorial office.*

## LITERATURE CITED

Abel, Andrew, and Frederic S. Mishkin. "An Integrated View of Tests of Rationality, Market Efficiency and the Short-Run Neutrality of Monetary Policy." *Journal of Monetary Economics* 11 (January 1983), 3–24.

Attfield, Clifford L. F., D. Demery, and Nigel W. Duck. "A Quarterly Model of Unanticipated Monetary Growth, Output and the Price Level in the U.K., 1963–1978." *Journal of Monetary Economics* 8 (November 1981), 331–50.

[10]The $\Delta_1$ parameter for the United States implies noninvertibility, indicating seasonal overdifferencing. Estimates with only first differences in the endogenous variable do not exhibit any significant influence of unexpected inflation, either, but are plagued with other problems.

[11]These findings are in agreement with Darby (1983).

[12]Examples are provided by Mankiw and Shapiro (1985) and Plosser and Schwert (1978) for the permanent income hypothesis and the relationship between money and income respectively.

Barro, Robert J. "Unanticipated Money, Output, and the Price Level in the United States." *Journal of Political Economy* 86 (August 1978), 549–80.

Barro, Robert J., and Mark Rush. "Unanticipated Money and Economic Activity." In *Rational Expectations and Economic Policy*, edited by Stanley Fischer, pp. 23–73. Chicago: University of Chicago Press, 1980.

Blejer, Mario I., and Leonardo Leiderman. "On the Real Effects of Inflation and Relative-Price Variability: Some Empirical Evidence." *Review of Economics and Statistics* 62 (November 1980), 539–44.

Box, George E. P., and Gwilym M. Jenkins. *Time Series Analysis, Forecasting and Control*, 2nd ed. San Francisco: Holden Day, 1976.

Chan, K. Hung, Jack C. Hayya, and J. Keith Ord. "A Note on Trend Removal Methods: The Case of Polynomial Regression versus Variate Differencing." *Econometrica* 45 (April 1977), 737–44.

Cuddington, John T. "Simultaneous-Equations Tests of the Natural Rate and Other Classical Hypotheses." *Journal of Political Economy* 88 (June 1980), 539–49.

Darby, Michael R. "Actual versus Unanticipated Changes in Aggregate Demand Variables: A Sensitivity Analysis of the Real-Income Equation." In *The International Transmission of Inflation*, edited by Michael R. Darby and James R. Lothian, pp. 273–88. Chicago: University of Chicago Press, 1983.

Evans, Paul. "The Effects on Output of Money Growth and Interest Rate Volatility in the United States." *Journal of Political Economy* 92 (April 1984), 204–22.

Fair, Ray C. "An Analysis of the Accuracy of Four Macroeconometric Models." *Journal of Political Economy* 87 (August 1979), 701–18.

Fuller, Wayne A. *Introduction to Statistical Time Series.* New York: Wiley, 1976.

Gordon, Robert J. "Price Inertia and Policy Ineffectiveness in the United States, 1890–1980." *Journal of Political Economy* 90 (December 1982), 1087–1117.

Granger, Clive W. J., and P. Newbold. "Spurious Regressions in Econometrics." *Journal of Econometrics* 2 (May 1974), 111–20.

Haraf, William S. "Test of a Rational Expectations-Structural Neutrality Model with Persistent Effects of Monetary Disturbances." *Journal of Monetary Economics* 11 (January 1983), 103–16.

Hasza, David P., and Wayne A. Fuller. "Testing for Nonstationary Parameter Specifications in Seasonal Time Series Models." *The Annals of Statistics* 10 (November 1982), 1209–16.

Hoffman, Dennis L., and Don E. Schlagenhauf. "An Econometric Investigation of the Monetary Neutrality and Rationality Propositions from an International Perspective." *Review of Economics and Statistics* 64 (November 1982), 562–71.

Hoffman, Dennis L., Stuart A. Low, and Don S. Schlagenhauf. "Tests of Rationality, Neutrality and Market Efficiency: A Monte Carlo Analysis of Alternative Test Statistics." *Journal of Monetary Economics* 14 (November 1984), 339–63.

Lawrence, Colin. "Rational Expectations, Supply Shocks and the Stability of the Inflation-Output Tradeoff, Some Time Series Evidence for the United Kingdom, 1957–1977." *Journal of Monetary Economics* 11 (March 1983), 225–45.

Leiderman, Leonardo. "Expectations and Output-Inflation Tradeoffs in a Fixed-Exchange-Rate Economy." *Journal of Political Economy* 87 (December 1979), 1285–1306.

Lucas, Robert E., Jr. "Some International Evidence on Output-Inflation Tradeoffs." *American Economic Review* 63 (June 1973), 326–34.

Makin, John H. "Anticipated Money, Inflation Uncertainty and Real Economic Activity." *Review of Economics and Statistics* 64 (February 1982), 126–34.

Mankiw, N. Gregory, and Matthew D. Shapiro. "Trends, Random Walks, and Tests of the Permanent Income Hypothesis." *Journal of Monetary Economics* 16 (September 1985), 165–74.

McCallum, Bennett T. "On "Real" and "Sticky-Price" Theories of the Business Cycle." *Journal of Money, Credit, and Banking* 18 (November 1986), 397–414.

Mishkin, Frederic S. "Does Anticipated Policy Matter? An Econometric Investigation." *Journal of Political Economy* 90 (February 1982), 22–51.

Nelson, Charles R., and Heejoon Kang. "Spurious Periodicity in Inappropriately Detrended Time Series." *Econometrica* 49 (May 1981), 741–51.

———. "Pitfalls in the Use of Time as an Explanatory Variable in Regression." *Journal of Business and Economic Statistics* 2 (January 1984), 73–82.

Nelson, Charles R., and Charles I. Plosser. "Trends and Random Walks in Macroeconomic Time Series." *Journal of Monetary Economics* 10 (September 1982), 139–62.

Pagan, Adrian. "Econometric Issues in the Analysis of Regressions with Generated Regressors." *International Economic Review* 25 (February 1984), 221–47.

Phillips, P. C. B. "Understanding Spurious Regressions in Econometrics." *Journal of Econometrics* 33 (December 1986), 311–40.

Plosser, Charles I., and G. William Schwert. "Money, Income, and Sunspots: Measuring Economic Relationships and the Effects of Differencing." *Journal of Monetary Economics* 4 (November 1978), 637–60.

Sargent, Thomas J. "A Classical Macroeconometric Model for the United States." *Journal of Political Economy* 84 (April 1976), 207–37.

Stulz, René M., and Walter Wasserfallen. "Macroeconomic Time-Series, Business Cycles and Macroeconomic Policies." *Carnegie-Rochester Conference Series on Public Policy* 22 (Spring 1985), pp. 9–53.

Taylor, John B. "Estimation and Control of a Macroeconomic Model with Rational Expectations." *Econometrica* 47 (September 1979), 1267–86.

Wasserfallen, Walter. "Forecasting, Rational Expectations and the Phillips-Curve, An Empirical Investigation." *Journal of Monetary Economics* 15 (January 1985), 7–27.

———. "Nonstationarities in Macroeconomic Time Series—Further Evidence and Implications." *Canadian Journal of Economics* 19 (August 1986), 498–510.