A Cautionary Note on the Order of Integration of Post-War Aggregate Wage, Price, and Productivity Measures

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Abstract

This paper investigates the order of integration of aggregate wage, price and productivity measures for the U.S.. Our investigation differs from previous studies as we employ recently developed tests that allow, under the alternative hypothesis, for structural change between periods in which the data is integrated of order zero, I(0), and integrated of order one, I(1). The tests reveal that some of the time series examined are neither exclusively I(0) or I(1), and that in fact breaks in the order of integration have occurred, suggesting the need for caution when undertaking Granger-causality tests involving these variables.

JEL classification: C22; E31

Keywords: Unit root; Structural breaks; Causality

1. Introduction

A number of previous studies have found conflicting evidence on the order of integration of aggregate measures of labor market pressures as well as measures of inflation for the U.S. over the post-war period. Whether or not such data is found to be integrated of order zero, I(0), or integrated of order one, I(1), appears to depend on the sample period examined. Such findings may be the result of a structural change in the data generating process of the variables in question. For some time the importance of allowing for structural change in the deterministic components of economic time series has been recognized, particularly when undertaking unit root tests, however, more recently it has been proposed that practitioners should also allow for the possibility of structural change in the persistence of economic time series. For example, Kim (2000) (Corrigendum, Kim et al., 2002), Busetti and Taylor (2001) and Leybourne *et al*. (2002) develop hypothesis tests that allow, under the alternative hypothesis, for breaks between I(0) and I(1) regimes, and illustrate the value of their tests with empirical applications to economic time series. If the economic variables mentioned above contain such structural change, this could be responsible for the conflicting evidence on the order of integration of the data from conventional unit root tests. The low and stable inflation which has existed since the mid-1980s, combined with the visibly changing nature of labor market variables such as unit labor cost and productivity measures, and the conflicting evidence on the order of integration of these variables from conventional unit root tests, suggests that such variables may have undergone a structural change between periods of I(0) and I(1) behavior. If this is true, it has important implications for studies that investigate causal linkages between such variables. Along these lines, He and Maekawa (2001) have shown that the *F*-statistic for testing Granger causality will find spurious causality (with high probability) between two independent variables where one or both of them is, or are, non-stationary.

The issue of whether wage inflation causes price inflation is a fundamental one for macroeconomists, however, the empirical evidence on this issue for the U.S. is mixed and also appears to depend on the sample period employed, leading us to suspect that there may have been changes in the order of integration of the wage and price data used. At the macro level, Mehra (1991) investigates the causal linkage between the aggregate price level and productivity adjusted wages. He finds that inflation and growth in unit labor costs are correlated in the long-run, and that inflation surprisingly Granger-causes growth in unit labor costs. Using unit root tests and the cointegration methodology, as well as alternative measures of the price level and different sample periods, Mehra (1993, 2000) examines the robustness of his earlier results, and demonstrates that aggregate unit labor costs and the aggregate price level contain a common stochastic trend. He finds bi-directional Granger-causality between these variables.[1]

More recently, Mehra (2000) demonstrates that (GDP deflator) inflation always helps to predict wage growth and this relationship is stable across all sample periods, but that wage growth predicts inflation only during the period 1966Q1-1993Q4, a period during which inflation steadily accelerated. Mehra (1991, 1993, 2000) questions the cost-push explanation of inflation. He shows that prices explain wages, but that wages are not a causal factor in determining inflation. Hu and Trehan (1995) and Gordon (1998) report evidence consistent with the results of Mehra, and show that wage growth has no predictive content for inflation, also rejecting the cost-push view. However, Emery and Chang (1996) and Hess (1999) demonstrate that these findings are sensitive to the sample period employed, whilst Hess and Schweitzer (2000) find evidence that wage inflation has very little predictive power for price inflation in the 1990s. They find that wage inflation, measured by labor compensations, wages, or unit labor costs, are not reliable predictors of price inflation. Using Granger causality tests, Ghali (1999), in

[1] More specifically, Mehra (1993) reports that over the period 1956:Q1-1992:Q4 wage inflation is found to cause price inflation only when one uses the CPI to measure prices. If one uses the more general GDP price deflator he finds strong evidence that prices granger-cause wages.

contrast to the findings of Mehra and others, finds that wage growth does help to predict inflation, supporting the cost-push view.

Rissman (1995) examines whether sectoral wage growth causes inflation. He finds that in most industries the direction of causality is from prices to wages. Only in the manufacturing and retail trade industries does productivity adjusted wage growth help to predict inflation. Clark (1997) employs in-sample and out-of-sample Granger-causality tests and finds little evidence that producer prices Granger-cause consumer prices, whilst Hogan (1998) finds that unit labor costs (in a Phillips curve equation framework) do not help to predict inflation. A related but also relatively unexplored relationship is that between real wages and productivity. Although this link is the building block of many macroeconomic models, and is frequently cited in intermediate macroeconomic textbooks (e.g. Gordon 2000), few empirical works have tested this relationship. In a competitive labor market deviations of real wages from productivity are expected to be stationary. This is equivalent to deviations of price from unit labor costs being stationary. [2]

The purpose of this paper is to re-examine the evidence on the order of integration of aggregate measures of labor market pressures as well as measures of inflation for the U.S. over the post-war period, using conventional unit root tests, and the recently developed tests of Kim (2000) and Leybourne *et al*. (2002); henceforth, LKSN. Kim (2000) develops a nonparametric residual based test that takes I(0) as the null and considers the alternative of either a structural break in the series from I(0) to I(1), or a break from I(1) to I(0). LKSN use regression-based augmented Dickey-Fuller (ADF) type statistics to test the null hypothesis that a time series is I(1) against the alternative of either a change from I(0) to I(1), or from I(1) to I(0).

In this paper the following series are considered; growth in unit labor costs, growth in wage compensation, inflation (measured by the personal consumption expenditure (PCE)

[2] Pioneering work investigating the linkage between real wages and productivity using aggregate data for the UK is Hall (1986, 1989). He finds a one-to-one linkage between real wages and productivity.

deflator, the consumer price index (CPI), the gross domestic product (GDP) price deflator), growth in adjusted monetary base, and growth in labor productivity. All variables are examined over the period 1959:Q1-1999:Q3. We also examine the behavior of deviations of real wages from labor productivity and deviations of the aggregate price level from productivity adjusted wages over the same period. From the tests employed in this paper we find evidence suggesting that a number of the variables examined are neither exclusively I(0) or I(1) over this period, and that in fact structural changes between I(0) and I(1) have occurred. In light of the recent evidence of He and Maekawa (2001), the implication is that practitioners should be extremely cautious when investigating the causal relationships between these variables. The paper is organized as follows. Section 2 discusses the different econometric procedures employed, Section 3 discusses our empirical results and Section 4 presents a summary and conclusion.

2. Econometric Methodology

2.1 Data and conventional unit root tests

The data employed in this paper is for the period 1959:Q1-1999:Q3 and is the same data that is examined by Hess and Schweitzer (2000).[3] The measure of money is the Federal Reserve Bank of St. Louis adjusted monetary base. Three price measures are used; the PCE deflator (seasonally adjusted, 1992=100), the CPI-U index (1982-84=100) and the GDP price deflator (1996=100). Productivity is measured as non-farm business productivity and is seasonally adjusted with 1992=100. Unit labor costs are non-farm business unit labor costs and are seasonally adjusted with 1992=100. Wages are non-farm business total compensation, are seasonally adjusted with 1992=100. All data are transformed by taking natural logs, and then computing annualized growth rates from the quarterly data. Initially, before investigating possible change in the order of integration of this data we apply a

[3] We would like to thank Mark Schweitzer for providing the data.

number of conventional unit root tests. Two of the most commonly used unit root tests in the literature are the augmented Dickey-Fuller (ADF) test of Said and Dickey (1984) and Phillips-Perron test (PP *Z*-tests) developed in Phillips and Perron (1988). It is well known however that the ADF and PP tests have low power against local stationary alternatives. Elliot, Rothenberg and Stock (1996) (ERS) developed a version of the ADF test that uses local GLS detrending to increase the power of the original test (DF-GLS). Another serious problem associated with the ADF and PP tests is that they suffer from serious size distortions when the data generating process (DGP) has negative moving average terms. Schwert (1987, 1989), Phillips and Perron (1988), Pantula (1991), Ng and Perron (1995, 2000) and Perron and Ng (1996) demonstrate that the empirical size of conventional ADF and PP tests approach unity as the sum of the MA parameters in a univariate process approach negative one. Perron and Ng (1996) extend the work of ERS by developing modified versions of the PP tests that have much better size properties than the conventional PP tests in the presence of moving average terms, but also retain the power of the DF-GLS test of ERS. These unit root tests also utilize the local GLS detrending method and are collectively referred to as *M*-tests (as with the conventional PP tests there are several types; specifically, $MZ(\alpha)$, $MZ(t)$, and the MSB test). The decrease in size and increase in power are enhanced when one chooses the lag length based on the modified information criteria (MIC) developed in Ng and Perron (2000). In our initial investigation of the above mentioned data, we employ the ADF, PP, DF-GLS, tests and the *M*-tests of Ng and Perron (2000). In all cases, lags are chosen using the MIC criteria of Ng and Perron (2000).

2. 2 Tests for change in the order of integration of a time series

Tests of the null hypothesis of level-stationarity, I(0), have recently been developed by Kim (2000) (Corrigendum, Kim et al., 2002)[4] that allow under the alternative hypothesis for breaks from I(0) to I(1) and I(1) to I(0). Specifically, Kim considers testing the following null hypothesis;

$$H_0 : y_t = r_0 + z_t \qquad t = 1,2,...,T\ ,$$

where r_0 is a constant and z_t is a stationary variable satisfying conventional regularity assumptions (see Kim 2000, Assumption 1), against the alternative hypothesis that y_t is a stationary process, with a break to a period of higher persistence (including I(0) to I(1)). The alternative hypothesis is written as;

$$H_1 : y_t = r_0 + z_{t,0} \qquad t = 1,2,..., \tau T$$

$$y_t = r_1 + z_{t,1} \qquad \tau T + 1,...,T$$

where τ is the break fraction ($0 < \tau < 1$), $z_{t,0}$ is a stationary process, $z_{t,1}$ is a process of higher persistence, and r_0 and r_1 are constants. Kim (2000) also considers testing the null hypothesis of level-stationarity against the alternative hypothesis of a break from a period of higher persistence to a period of lower persistence (including I(1) to I(0)), and the null hypothesis of trend-stationarity against the alternative hypotheses of a break from lower persistence to higher persistence and *vice-versa*.

Kim's test procedure has two steps; first test H_0 against the alternative hypothesis of breaks in persistence; then, if the null hypothesis is rejected, estimate the break-point. The test statistic proposed by Kim (2000) for testing H_0 against breaks in persistence is

[4] Henceforth we will refer only to Kim (2000), but in all empirical work we take into account the correction of Kim et al. (2002). The problem with the original Kim (2000) tests was identified by Stephen Leybourne and is also discussed in Busetti and Taylor (2001).

$$\Xi_T(\tau) = \frac{[(1-\tau)T]^{-2}\sum_{\tau T+1}^{T} S_{1,t}(\tau)^2}{(\tau T)^2 \sum_{1}^{\tau T} S_{0,t}(\tau)^2}, \tag{1}$$

where $S_{0,t}(\tau) = \sum_{i=1}^{t} \tilde{z}_{0,t}(\tau)$ for $t = 1,\ldots,\tau T$ and $S_{1,t}(\tau) = \sum_{i=\tau T+1}^{t} \tilde{z}_{1,t}(\tau)$ for $t = \tau T + 1,\ldots,T$ with $\tilde{z}_{0,t}(\tau)$ being the residuals from the regression of y_t on an intercept (or an intercept and trend if the null is trend-stationarity) for $t = 1,2,\ldots,\tau T$, and $\tilde{z}_{1,t}(\tau)$ being the residuals from the regression of y_t on an intercept (or an intercept and trend if the null is trend-stationarity) for $t = \tau T + 1,\ldots,T$. Since the point at which a break occurs (τT) is not known, for an operational version of the test, Kim proposes combining $\Xi_T(\tau)$ with some established procedures. The three procedures employed are the maximum Chow (MAX CHOW) approach of Davies (1977),

$$\Xi_T(\tau) \equiv \max_{\tau\in\theta} \Xi_T(\tau) \tag{2}$$

where θ is a compact subset of $(0,1)$; the mean score approach of Hansen (1991)

$$\Xi_T(\tau) \equiv \int_{\tau\in\theta} \Xi_T(\tau)\,d\tau, \tag{3}$$

and the mean-exponential (MEAN-EXP) approach of Andrews and Ploberger (1994)

$$\log E\exp[\Xi_T(\tau)] \equiv \log\left\{\int_{\tau\in\theta} \exp[\Xi_T(\tau)]\,d\tau\right\}. \tag{4}$$

Kim derives the asymptotic distribution of $\Xi_T(\tau)$ for various null hypotheses and presents small and large sample critical values for the test statistics. Applying these tests to time series on U.S. inflation and the U.S. federal government budget deficit series over the post-war period, Kim finds significant evidence of structural change from I(0) to I(1). For those series for which a rejection of the unit root null hypothesis is obtained, the second step of the testing procedure proposed by Kim (2000) involves estimating the break-point. This can be done using the statistic

$$\Lambda_T(\tau) = \frac{\sum_{\tau T+1}^{T} \tilde{z}_{1,t}(\tau)^2 / [(1-\tau)T]^2}{\sum_{1}^{\tau T} \tilde{z}_{0,t}(\tau)^2 / [\tau T]}$$

where $\tilde{z}_{0,t}(\tau)$ and $\tilde{z}_{1,t}(\tau)$ are defined as before. A consistent estimator of the break-point is given by searching over τ and choosing the value that yields $\text{argmax}\Lambda_T(\tau)$ when the null hypothesis is level-stationarity, and $\text{argmin}\Lambda_T(\tau)$ when the null hypothesis is trend-stationarity.

While the tests developed by Kim (2000) are undoubtedly a useful addition to the econometrician’s toolkit, they are tests against the null hypothesis of stationarity. Given the importance of the unit root hypothesis in applied econometrics it is not surprising that tests for change in the persistence of time series have been developed against the null hypothesis of a unit root. Such tests have recently been developed by LKSN. Consider the following model

$$y_t = d_t + \mu_t \qquad (4)$$

$$u_t = \alpha u_{t-1} + \phi(L)\Delta u_{t-1} + \varepsilon_t$$

where $d_t = z_t'\beta$, $\phi(L)$ is a lag polynomial of order $p-1$ with the roots of $1-\phi(L)$ being outside the unit circle, $z_t = [1, t]'$, $\beta = [\beta_0, \beta_1]'$ and ε_t is a martingale difference sequence satisfying conventional regularity assumptions (see LKSN, Assumption 1). The null hypothesis that LKSN consider is that y_t contains a unit root; that is, $\alpha = 1$. They focus on testing this null against two alternative hypotheses; a structural break from I(0) to I(1), and a structural break from I(1) to I(0).

Consider LKSN's tests of the null hypothesis of a unit root against the alternative hypothesis of a break from I(0) to I(1). All the LKSN tests exploit the GLS de-trending method of ERS, and the calculation of the tests begins by obtaining the GLS transformed data;

$$y_{\bar{\alpha}}(\tau) = [y_1, y_2 - \bar{\alpha}y_1, ..., y_{[\tau T]} - \bar{\alpha}y_{[\tau T]-1}]'$$

$$Z_{\overline{\alpha}}(\tau) = [z_1, z_2 - \overline{\alpha} z_1, ..., z_{[\tau T]} - \overline{\alpha} z_{[\tau T]-1}]',$$

where $\overline{\alpha} = 1 + \overline{c}/T$ for some $\overline{c} < 0$, $\tau \in (0,1)$ and T is the sample size. The next stage is to obtain the recursive GLS estimator of β by regressing the GLS transformed data $y_{\overline{\alpha}}(\tau)$ on $Z_{\overline{\alpha}}(\tau)$ and constructing the residual series $y_t^d = y_t - \hat{\beta}_0(\tau) - \hat{\beta}_1(\tau)t$. LKSN's test statistic for testing the null of a unit root against the alternative of a break from I(0) to I(1) is the infimum of the recursive *t*-statistic for testing $\rho(\tau) = 0$ in the model;

$$\Delta y_t^d = \rho(\tau) y_{t-1}^d + \sum_{j=1}^{p-1} \phi_j(\tau) \Delta y_{t-j} + \varepsilon_t \qquad t = 1,2,...,\tau T \tag{5}$$

Thus, the test statistic is, $DF_G^{f\,\inf} = \inf_{\tau \in \Lambda} DF_G^f(\tau)$. LKSN show that under the alternative hypothesis the value of τ at the infimum is a consistent estimator of the break-point.

The test statistic $DF_G^{f\,\inf}$ does not exploit the full information in the sample, using only those observations up to τT, however, LKSN also propose a more powerful version of this test statistic based on the full sample of data. The specification for calculating this statistic is

$$\Delta y_t^d = \overline{\rho}(\tau) D_t(\tau) y_{t-1}^d + \sum_{j=1}^{p-1} \phi_j(\tau) \Delta y_{t-j} + \hat{\varepsilon}_t \qquad t = 1,2,...,T \tag{6}$$

where $D_t(\tau)$ is a dummy variable;

$$D_t(\tau) = 1[t \leq \tau T], \tag{14}$$

and $\Delta y_t^d = \Delta y_t - \Delta \overline{y}_{(2)}$, $\Delta \overline{y}_{(2)} = (T - \tau T)^{-1} \sum_{s=\tau T+1}^{T} \Delta y_s$ when $t > \tau T$, with Δy_t^d being defined as in (13) for $t = 1,2,...,\tau T$. In this case the test statistic is the infimum of the *t*-statistic for testing $\overline{\rho}(\tau) = 0$ in (13);

$$\overline{DF}_G^{f\,\inf} = \inf_{\tau \in \Lambda} \overline{DF}_G^f(\tau) \tag{7}$$

To test the null hypothesis of a unit root against the alternative hypothesis of a break from I(1) to I(0), LKSN exploit the fact that a break from I(1) to I(0) corresponds to a break from I(0) to I(1) in the original series. They define $\tilde{y}_t = y_{T-t+1}$ and propose using the GLS transformed time reversed data $\tilde{y}_t$ and the following specification;

$$\Delta\tilde{y}_t^d = \tilde{\rho}(\tau)\tilde{y}_{t-1}^d + \sum_{j=1}^{p-1}\tilde{\phi}_j(\tau)\Delta\tilde{y}_{t-j}^d + \tilde{\eta}_t \qquad t = 1,2,\ldots,(1-\tau)T \, . \tag{8}$$

The test statistic is

$$DF_G^{r\inf} = \inf_{\tau\in\Lambda} DF_G^r(\tau)\,, \tag{9}$$

where $DF_G^{r\inf}$ is the *t*-statistic for testing $\tilde{\rho}(\tau) = 0$ in (8). A further test employing the full sample of time reversed data is also proposed;

$$\overline{DF}_G^{r\inf} = \inf_{\tau\in\Lambda} \overline{DF}_G^r(\tau)\,, \tag{10}$$

where $\overline{DF}_G^r(\tau)$ is the *t*-statistic for testing $\overline{\rho}(\tau) = 0$ in

$$\Delta\tilde{y}_t^d = \overline{\rho}(\tau)D_t(\tau)\tilde{y}_{t-1}^d + \sum_{j=1}^{p-1}\overline{\phi}_j(\tau)\Delta\tilde{y}_{t-j}^d + \overline{\eta}_t \qquad t = 1,2,\ldots,T \, . \tag{11}$$

3. Empirical Results

Our analysis focuses on the growth rate of seven variables and two common macro linkages. The seven variables considered are; the growth rate of per-unit labor costs (gulc), the growth rate of wage compensation (gcomp), the inflation rate measured using three different price indexes (the PCE deflator (gpce_p), the CPI (gcpi_p), and the GDP price deflator (ggdp_p)), the growth rate of the adjusted monetary base (gamb), and the growth rate of labor productivity (gy_n). The two linkages we examine are ln(price/per-unit labor costs), where price is measured by the gpce_p, gcpi_p and ggdp_p, and ln(real wage/prod), where prod denotes labor productivity and the real wage is computed using the above three price deflators. When the GDP price deflator is used we refer to real wages as rw1; rw2 refers to the use of the PCE in computing real wages, and rw3 refers to the use of the CPI in computing real wages.

Figure 1 and 2 plot the above mentioned series. Figure 1 plots the growth rates while Figure 2 plots the other series. From Figure 1, visually, it appears that there is a distinct change in the data generating process of the gulc, gcomp, gpce_p, gcpi_p and ggdp_p series, occurring in the early 1980s. Figure 2 shows, as expected that ln(price/ulc) and ln(real wage/prod) are inversely related. We next investigate these series by conducting the conventional unit root tests and the tests for change in the order of integration.

Table 1 and 2 report unit root tests for the above series, while Table 3 and 4 report tests for a change in the order of integration of the series.[5] We perform these tests including a deterministic trend and also without a trend. For all the test statistics, in each cell of the tables the case without a trend is reported first in the cell, and then the number below it corresponds to the test statistic for the case when a trend is included. For the LKSN tests we imposed the same lag-length as in the conventional unit root tests (thus determined assuming no break). We then conducted a sensitivity analysis to small changes in the lag-length (± 2 lags). As expected, such changes in lag-length do sometimes have an impact on the value of the calculated test statistic and the estimated break-date from the LKSN test, however, in all cases the impact is small and does not alter the conclusions reported below.

The first panel of Table 1 and 2 reports results for GLS detrending, while the lower panel reports OLS detrending results. From Table 1, the tests show that using GLS detrending, the unit root null cannot be rejected for gcomp, gpce_p, and ggdp_p. The PP test rejects the unit root null for gulc, gamb and gy_n. The DF-GLS test rejects the unit root null for gcpi_p when a trend is included, while all other tests find that the unit root null cannot be rejected for gcpi_p. All tests with the exception of the PP test find that the unit root null cannot be rejected for gulc and gy_n. When a trend is included, the *M*-tests and the DF-GLS test rejects the unit root null for gamb. Using OLS detrending, the $MZ(\alpha)$ and ADF test find that the unit root null cannot be

[5] We would like to thank Jae-Young Kim and Stephen Leybourne for providing the GAUSS code for Kim and LSKN tests.

rejected for gulc, gcomp, gpce_p, gcpi_p, and ggdp_p. The $Z(\alpha)$, $MZ(\alpha)$ and the ADF test reject the unit root null under OLS detrending for gamb and gy_n. The $Z(\alpha)$ test, under OLS detrending, rejects the unit root null for gulc, gcomp, and gcpi_p.

Overall, from these tests there is evidence suggesting that five of the seven growth rate series examined are I(0). The results of Table 2 indicate that, with one exception of rw1 under GLS detrending using the DF-GLS test and a trend included, the ln(price/ulc) and ln(real wage/ulc) series are all found to be I(1), regardless of which price index is used.

Table 3 reports the Kim and the LKSN tests for change in order of integration. In analyzing the Kim tests there are two alternative hypotheses. If the test statistic is less than the critical value in the left portion of the distribution then one rejects the I(0) null in favor of the alternative of a switch from I(1) to I(0). If the test statistic is greater than the critical value in the right portion of the distribution then the I(0) null is rejected in favor of the alternative of I(0) to I(1). The critical values used to evaluate the Kim tests are those given in Busetti and Taylor (2001), Table 2.1 for the case of no trend and Table 6.1 allowing for a trend. These critical values only cover the right portion of the distribution, thus to evaluate the left portion of the distribution we use improved critical values provided to us by Jae-Young Kim. Just as in Tables 1 and 2, for all the test statistics, in each cell of the tables, the case without a trend is reported first in the cell, and then the number below it corresponds to the test statistic for the case when a trend is included.

For gulc and gcomp (when a trend is omitted), two out of three of the Kim tests find that one can reject the null of stationarity in favor of the alternative of a switch from I(0) to I(1) occurring in 1990. When a trend is allowed for, two tests reject the null of stationarity for gulc, and three tests reject the null for gcomp. For the series gpce_p, gcpi_p, and gdp_p the Kim tests find a structural break in 1967 and reject the null of I(0) in favor of the alternative of a break from I(0) to I(1). For the gamb and gy_n series, all three tests (when no trend is included) reject the I(0) null in favor of the alternative of a break from I(1) to I(0) occurring in 1990.

With respect to the LKSN tests, we employ the sequential version of the test which utilizes the full sample of data. LKSN advocate doing the test on the data ordered both forward (f) in time and reverse (r) in time. The minimum test statistic from these two tests should then be compared with the appropriate critical value to determine if the null of a unit root can be rejected. In our tables, (f) refers to the test being employed on the data ordered naturally (forward in time) and therefore the alternative is I(0) to I(1), while (r) refers to the test on the reversed data and therefore the alternative hypothesis is I(1) to I(0). Again, the first number in the cell is for the case when no trend is included while the second number is for the case when a trend is allowed.

For the series gulc, gpce, and gy_n, we find that the null of a unit root cannot be rejected regardless of whether a trend is included or not. When no trend is included gcomp is found to contain a unit root, while the unit root null is rejected when a time trend is included in favor of the alternative of I(1) to I(0) with a break found in 1971. When no trend is included, the null of unit root is rejected for gcpi_p in favor of the alternative of switch from I(0) to I(1) occurring in 1981, however, the unit root null cannot be rejected for gcpi_p when a time trend is included. For ggdp_p, the null of a unit root cannot be rejected when no trend is included, but is rejected in favor of the alternative of a switch from I(0) to I(1) with a break occurring in 1980 when a trend is included. For the series gamb, the null of unit root is rejected regardless of whether a trend is included or not. For gamb, when a trend is not included we find in favor of the alternative of a switch from I(0) to I(1) with a break in 1989, and a break in 1987 when a trend is included.

From Table 4 we note that (with one exception) all three Kim tests find that for ln(cpi_p/ulc) and ln(rw3/prod) the null of I(0) is rejected in favor of a change from I(0) to I(1) occurring in 1981. This should not be surprising as both series use the CPI in computing this relation. Similarly, all three of Kim's tests find that for ln(pce/ulc) and ln(rw2/prod) the null of I(0) is rejected in favor of the alternative of a change from I(0) to I(1) occurring in 1973. For the series ln(ggdp_p/ulc) only the MAX-CHOW and the MEAN-EXP tests with a trend included reject the I(0) null in favor of a change from I(0) to I(1) occurring in 1981. Similar results are found for ln(rw1/prod) in that only the MAX-CHOW and MEAN-EXP tests with a trend included reject the I(0) null in favor of the alternative of a change from I(0) to I(1) occurring in 1981.

Interestingly, for all of the ln(price/ulc) and ln(real wage/prod) series examined the LKSN test fails to reject the null of a unit root. Recall, the null hypothesis of the Kim tests is that the data is I(0) with no breaks. Thus the absence of evidence of structural change from the LKSN tests suggests that some of the rejections of the null of I(0) from the Kim tests could be because the data is in fact I(1) with no break.

The results from the Kim and LKSN tests are difficult to compare as they are testing two different null hypotheses. Furthermore, the LKSN tests are specifically designed to allow for breaks between I(0) and I(1) under the alternative, while the Kim tests are designed to allow for breaks in *persistence*, which includes breaks between periods of I(0) behaviour with varying degrees of persistence. Thus, one should not expect *a priori* that the Kim tests and LKSN tests will choose the same types of breaks or break-points. In order therefore to clarify the results of these tests, it is advisable to apply conventional unit root tests to the sub-samples given by the estimated break-points.[6] In Tables 5 and 6, we report the results from applying the appropriate conventional unit root test (appropriate in terms of whether or not a trend is included) to a selection of the growth rate data for the sub-samples given by the break-points of the LKSN and Kim tests. Only those series for which a rejection of the respective null hypotheses are considered, and only series with a break-point in 1980 or later are considered (to avoid having

[6] We would like to thank an anonymous referee for drawing our attention to these points.

particularly small sub-samples). In all cases one or two observations either side of the estimated break-point are left out.

Consider first Table 5 which contains the results for the LKSN break-points. For gcpi_p the results do not lend support to the conclusion from the application of the LKSN tests, for ggdp_p and gamb however, there is strong support for the finding of a break from trend-stationarity to I(1) – exactly what the LKSN tests suggest. Interestingly, from Table 6 which contains the results for the break-point of 1990 as chosen by the Kim tests, there is much less support for a break in any of the series. Indeed, since there are so few rejections of the unit root null, the evidence in this table suggests that in a number of instances the rejections by the Kim test reported in Table 3 may simply be due to the series being I(1) with no change in persistence (this would support the fact that for a number of these series the LKSN test does not yield a rejection of the unit root null).

He and Maekawa (2001) have illustrated the importance of ensuring that Granger-causality tests are carried out using data that is I(0). In light of this, our evidence of possible breaks between I(0) and I(1) (particularly for the ggdp_p and gamb series) suggests the need for caution when interpreting the results of Granger-causality tests applied to this data. To reinforce this point we present an illustrative example concerning the short-run dynamics of wages and prices.[7] Table 7 reports the *F*-statistics for testing Granger-causality between the growth rate of wages and growth rate of prices. These were calculated from a regression including four lags of the growth rate of wages, growth rate of prices, and growth rate of money supply. We undertake the test for two different sub-samples of data; the period 1959:1 – 1979:4 and the period 1971:4 – 1979:4. Wage growth is measured by gcomp, price growth by ggdp_p, and money supply growth by gamb. The models were estimated by OLS and we carried out the test under two different assumptions regarding the order of integration of the series. The second column of the

[7] We would like to thank an anonymous referee for suggesting this addition to the paper.

table contains results assuming that all the data is I(0) (and a deterministic trend is included in the models). The third column contains results assuming that all the data is I(1) and is differenced once before carrying out the Granger-causality test.

Consider the results for the longer sample period for which the data is assumed to be I(0). These results suggest that prices are caused by wages, but that the converse is not true. However if the data is assumed to be I(1) and first-differenced, no causality is found. Recall that the LKSN test finds that over the period 1959:1 – 1979:4, ggdp_p and gamb are both trend-stationary but that gcomp begins as an I(1) process with a break to I(0) in 1971. Therefore according to the LKSN test, none of the results for the longer sample can be trusted since they are obtained using data that, for part of the sample, is of a different order of integration. Consider the results for the shorter sample period 1971:4 – 1979:4. Here, according to the LKSN test the results from the model that assumes all the data is I(0) can be trusted - in as much as the order of integration assumed is correct. In this case however there is no statistically significant evidence of causality from wages to prices, but there is now statistically significant evidence of causality from prices to wages! These results suggests that our finding of causality from wages to prices using the sample 1959:1 – 1979:4 may be driven by a break from I(1) to I(0) in the measure of wage growth used (gcomp), and that for the latter part of this period (1971:4 – 1979:4) causality is in fact from prices to wages. Of course, we recognize the limitations of this brief analysis of causality, however, this simple example is illustrative of just how sensitive Granger-causality tests involving these variables can be to the order of integration assumed and the sample period examined – reinforcing the main point of this paper that practitioners should be cautious when interpreting the results of such tests applied to the type of data examined here.

4. Conclusion

Previous empirical research has produced conflicting evidence on the order of integration of aggregate measures of labor market pressure and inflation in the U.S. over the post-war period. Similarly, previous empirical evidence on the causal relationship between wages and prices is conflicting, and appears to depend to some extent on the sample period examined. This suggests that the variables examined may have undergone structural change between I(0) and I(1) regimes. This paper has investigated the order of integration of a number of wage, price and productivity measures for the U.S. over the post-war period, and two common macro linkages. The analysis differs from previous work with this data as, in addition to conventional unit roots tests, we have applied two recently developed tests that under the alternative hypothesis allow for changes in the order of integration of the time series; that is, changes from either I(0) to I(1), or I(1) to I(0). From the conventional unit root tests employed, there is evidence against the null hypothesis of a unit root for five of the seven growth rate series examined. Thus, there is evidence from the conventional tests suggesting that a number of these time series may be I(0). However, from the tests developed by Kim (2000) of the null hypothesis of I(0), allowing under the alternative hypothesis for breaks between I(0) and I(1), for the growth rate series there is evidence against the null in favor of change in the order of integration for all of the seven series, with breaks being estimated to have occurred in either 1967 or 1990. From the tests developed by LKSN of the null hypothesis that the time series is I(1), also allowing for breaks between I(0) and I(1) under the alternative hypothesis, for the growth rate data there is evidence against the null in favor of breaks between I(0) and I(1) for four of the seven series, with breaks being estimated to have occurred in the early 1970s, early 1980s and late 1980s.

We have also examined some key macroeconomic linkages, namely the deviations of real wages from labor productivity and the aggregate price level and productivity adjusted wages over the same period. In a competitive labor market deviations of real wages from productivity are expected to be stationary. This is equivalent to deviations of price from unit labor costs being stationary. The conventional unit root tests generally do not reject the null of I(1); the Kim tests reject the null of I(0) in favor of a break from I(0) to I(1), whilst the LKSN test fails to reject the null of I(1). These results suggest that these linkages could be I(1) with no break, and thus the Kim tests are rejecting the null of I(0) in favor of a break between I(0) and I(1), when in fact no break has occurred.

Since the LKSN test may yield rejections of the null of I(1) if the data is I(0) with no break, and that the Kim tests may yield rejections of the null of I(0) if the data is in fact I(1) with no break, it is important that we have applied both tests in our investigation. Furthermore, for the growth rate series we also investigate changes in the order of integration in more detail by applying conventional unit root tests to sub-samples of data using the break-dates suggested by the Kim and LKSN tests. This investigation reveals support for the results of the LKSN tests, which for example suggest that breaks in the inflation and money supply series occurred in the 1980s. However, there is less support for the breaks suggested by the Kim tests. Given the propensity of evidence in this paper suggesting the possibility of breaks in the order of integration of the series examined, and the evidence of He and Maekawa (2001) on the possibility of spurious results if Granger-causality tests are applied to I(1) data, we warn that practitioners should be cautious when interpreting the results of Granger-causality tests involving these variables.

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Table 1
Conventional unit root tests for growth rates

GLS detrending	gulc	Gcomp	gpce_p	gcpi_p	ggdp_p	gamb	gy_n
$Z(\alpha)$	-16.781** -26.051**	-8.001 -11.783	-4.446 -6.042	-6.859 -11.088	-3.616 -4.959	-9.737** -31.334**	-35.255** -77.458**
DF-GLS	-1.489 -1.852	-1.094 -1.302	-1.431 -1.589	-1.740 -2.135**	-1.300 -1.437	-1.966 -3.468**	-1.260 -2.680
$MZ(\alpha)$	-3.578 -4.845	-2.241 -2.919	-4.077 -5.369	-6.271 -9.814	-3.388 -4.501	-6.578 -18.274**	-1.770 -4.395
$MZ(t)$	-1.332 -1.493	-1.044 -1.199	-1.428 -1.584	-1.765 -2.187	-1.299 -1.414	-1.714 -3.022**	-0.940 -1.397
OLS detrending							
$Z(\alpha)$	-34.242** -34.707**	-17.061** -19.256	-7.914 -8.012	-14.829** -14.778	-7.267 -7.721	-29.661** -36.891**	-154.900*** -167.131***
ADF	-2.317 -2.382	-1.680 -1.965	-1.848 -1.870	-2.553 -2.540	-1.807 -1.913	-3.962** -4.019**	-6.141** -6.289**
$MZ(\alpha)$	-6.683 -6.893	-3.779 -4.179	-6.454 -6.493	-12.280 -12.273	-5.843 -6.047	-17.056** -21.299**	-66.046** -76.211**

Notes: **, and *** denote significance at the 5% and 1% levels respectively.

Table 2
Conventional unit root tests for macro linkages

GLS detrending	ln(gdp_p/ulc)	ln(pce_p/ulc)	ln(cpi_p/ulc)	ln(rw1/prod)	ln(rw2/prod)	ln(rw3/prod)
$Z(\alpha)$	0.367 -15.638	0.743 -6.940	1.734 -5.412	0.365 -16.552	0.739 -7.508	1.725 -5.960
DF-GLS	0.2501 -2.864	0.489 -1.866	2.412 -1.624	0.249 -2.950**	0.486 -1.945	2.378 -1.708
$MZ(\alpha)$	0.4168 -14.878	0.786 -6.777	1.788 -5.249	0.416 -15.698	0.784 -7.324	1.779 -5.786
$MZ(t)$	0.2841 -2.725	0.517 -1.822	2.486 -1.574	0.285 -2.797	0.515 -1.897	2.453 -1.658
OLS detrending						
$Z(\alpha)$	-2.713 -15.666	-1.129 -7.618	0.3423 -9.867	-2.987 -16.848	-1.307 -7.928	0.286 -9.872
ADF	-1.225 -2.866	-0.579 -1.992	0.3536 -2.462	-1.311 -2.983	-0.652 -2.019	0.289 -2.418
$MZ(\alpha)$	-1.370 -14.900	-0.431 -7.228	1.150 -7.914	-1.492 -15.918	-0.515 -7.610	1.140 -8.206

Notes: ** denotes significance at the 5% level. rw1 is the real wage employing the GDP price deflator, rw2 employs the PCE price deflator, and rw3 employs the CPI.

Table 3
Kim and LKSN tests for growth rates

Kim tests	gulc	gcomp	gpce_p	gcpi_p	ggdp_p	gamb	gy_n
MAX CHOW	25.42** 3.675	44.82*** 15.12***	77.41*** 38.53***	47.16*** 46.66***	36.09*** 33.89***	.4289** 3.984	1.399* 1.641
MEAN SCORE	1.356 0.7363	2.074 3.087**	3.063 2.079	2.104 2.897**	1.871 2.289	.1102** 0.9172	.1965** 0.3781
MEAN-EXP	20.34** 1.328	10.64*** 39.74***	72.41*** 33.72***	42.15*** 41.86***	31.18*** 29.26***	-.3135** 1.422	-.1342** 0.1688
Break year	1990 1990	1990 1990	1967 1967	1967 1967	1967 1967	1990 1990	1990 1990
LKSN tests							
Minimum of (f) and (r) tests	-1.931 (f) -2.804 (f)	-2.687 (f) -3.842** (r)	-2.343 (r) -3.261 (f)	-3.261** (r) -3.242 (r)	-1.842 (r) -3.687* (f)	-3.870** (f) -5.013** (f)	-2.346 (f) -2.421 (f)
Break year	NA NA	NA 1971	NA NA	1981 NA	NA 1980	1989 1987	NA NA

Notes: *, **, and *** denote significance at the 10%, 5% and 1% levels respectively. The critical values used to evaluate the Kim tests outside of the bold border are those given in Busetti and Taylor (2001), Table 2.1 for the case of no trend and Table 6.1 allowing for a trend. The cells contained within the bold border relate to rejections from Kim's tests in favor of a break from I(1) to I(0) – the critical values used to evaluate these tests were provided to us by Jae-Young Kim. The rejections in any of the other cells associated with Kim's test are rejections in favor of a break from I(0) to I(1). (f) refers to the forward test $\overline{DF}_G^{f\,\inf} = \inf_{\tau\in\Lambda} \overline{DF}_G^{f}(\tau)$, (r) refers to the reverse test $\overline{DF}_G^{r\,\inf} = \inf_{\tau\in\Lambda} \overline{DF}_G^{r}(\tau)$. Critical values for the LKSN tests can be found in LKSN Table 2.

Table 4
Kim and LKSN tests for macro linkages

Kim tests	ln(gdp_p/ulc)	ln(pce_p/ulc)	ln(cpi_p/ulc)	ln(rw1/prod)	ln(rw2/prod)	ln(rw3/prod)
MAX CHOW	7.710 11.43**	41.68*** 46.76***	77.99*** 15.35***	7.855 11.88**	44.05*** 49.54***	79.19*** 16.11***
MEAN SCORE	1.509 1.081	13.27*** 2.107**	12.61*** .8850	1.555 1.083	13.97* 2.117	13.07 .8858
MEAN- EXP	4.269 7.627**	37.88*** 42.10***	72.91*** 10.95***	4.467 7.940**	40.17*** 44.73***	74.12*** 11.66***
Break year	1981 1981	1981 1981	1973 1973	1981 1981	1982 1982	1973 1973
LKSN tests						
Minimum of (f) and (r) tests	-1.808 (f) -2.849 (f)	-2.668 (f) -2.994 (f)	-1.055 (f) -3.137 (r)	-1.884 (f) -2.937 (f)	-2.756 (f) -3.043 (f)	-1.129 (f) -3.260 (r)
Break year	NA NA	NA NA	NA NA	NA NA	NA NA	NA NA

Notes: rw1 is the real wage employing the GDP price deflator, rw2 employs the PCE price deflator, and rw3 employs the CPI. *, **, and *** denote significance at the 10%, 5% and 1% levels respectively. The critical values used to evaluate the Kim tests are those give in Busetti and Taylor (2001), Table 2.1 for the case of no trend and Table 6.1 allowing for a trend.. (f) refers to the forward test $\overline{DF}_G^{f\,\inf} = \inf_{\tau\in\Lambda} \overline{DF}_G^{f}(\tau)$, (r) refers to the reverse test $\overline{DF}_G^{r\,\inf} = \inf_{\tau\in\Lambda} \overline{DF}_G^{r}(\tau)$. Critical values for the LKSN tests can be found in LKSN Table 2.

Table 5
Conventional unit root tests on sub-samples of a selection of growth rates: LKSN choice of break

GLS detrending	gcpi_p 1959:1 - 1981:2	gcpi_p 1982:1 - 1999:3	ggdp_p 1959:1 - 1980:2	ggdp_p 1981:1 - 1999:3	gamb 1959:1 - 1987:2	gamb 1988:1 - 1999:3
$Z(\alpha)$	-3.254	-4.828	-25.443**	-6.217	-48.436***	-10.343
DF-GLS	-.905	-.417	-3.063**	-1.395	-4.202***	-2.257
$MZ(\alpha)$	-3.004	-.951	-3.021	-3.909	-22.886**	-8.858
$MZ(t)$	-1.087	-.584	-18.779**	-1.374	-3.295**	-2.072
OLS detrending						
$Z(\alpha)$	-6.665	-15.960**	-26.405**	-21.653	-57.578***	-10.357
ADF	-1.644	-1.509	-3.089*	-2.810	-4.721***	-2.248
$MZ(\alpha)$	-5.508	-3.338	-19.467*	-8.161	-29.801**	-8.871

Notes: *, **, and *** denote significance at the 10%, 5% and 1% levels respectively.

Table 6
Conventional unit root tests on sub-samples of a selection of growth rates: Kim choice of break 1990

GLS detrending	gulc: 1959:1 - 1990:2	gulc 1991:1 - 1999:3	gcomp 1959:1 - 1990:2	gcomp 1991:1 - 1999:3	gamb 1959:1 - 1990:2	gamb 1991:1 - 1999:3	gy_n 1959:1 - 1990:2	gy_n 1991:1 - 1999:3
$Z(\alpha)$	-13.185**	-32.602***	-6.018	-9.427**	-8.790**	-7.240	-34.615**	-21.786
DF-GLS	-1.355	-1.851	-.945	-1.904	-1.847	-1.996	-1.071	-1.411
$MZ(\alpha)$	-2.970	-1.397	-1.518	-2.529	-4.800	-6.421	-1.170	.278
$MZ(t)$	-1.158	-.744	-.818	-1.096	-1.352	-1.770	-.657	.328
OLS detrending								
$Z(\alpha)$	-27.106**	-33.316**	-15.851**	-10.037	-24.512	-7.257	-78.160**	-21.826**
ADF	-2.269	-1.803	-1.713	-2.123	-3.704**	-1.992	-3.093*	-1.325
$MZ(\alpha)$	-6.748	-1.743	-3.328	-2.595	-12.032	-6.430	-10.486	.312

Notes: ** denotes significance at the 5% level.

Table 7
Granger test for causality between wages and prices for sub-samples of growth rates

Order of integration assumed	I(0)	I(1)
Sample period	1959:1 – 1979:4	
Prices caused by wages	2.145*	.859
Wages caused by prices	1.072	.390
Sample period	1971:4 – 1979:4	
Prices caused by wages	.843	1.489
Wages caused by prices	2.074*	.389

Notes: * denotes significance at the 10% level.

Figure 1

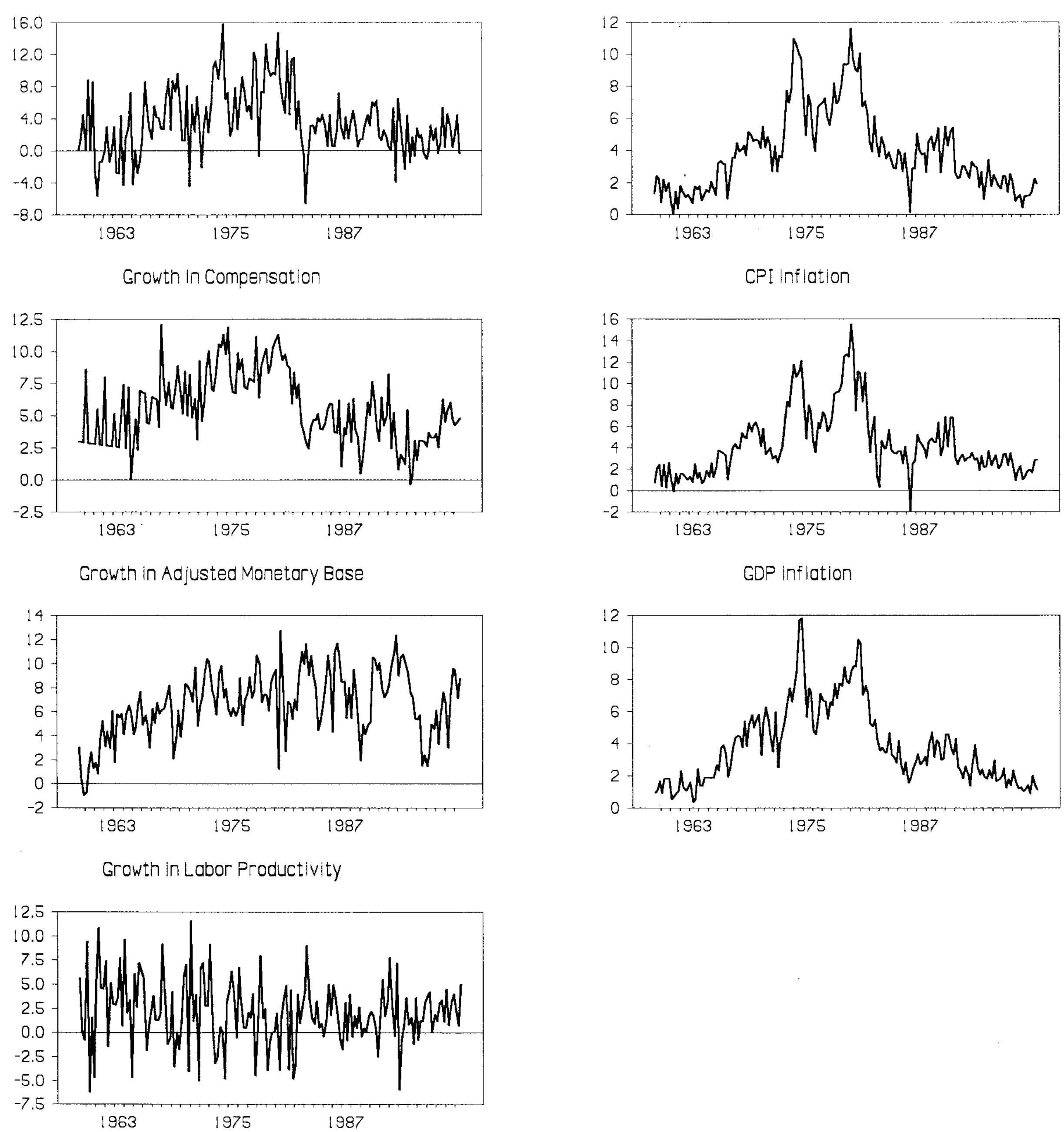
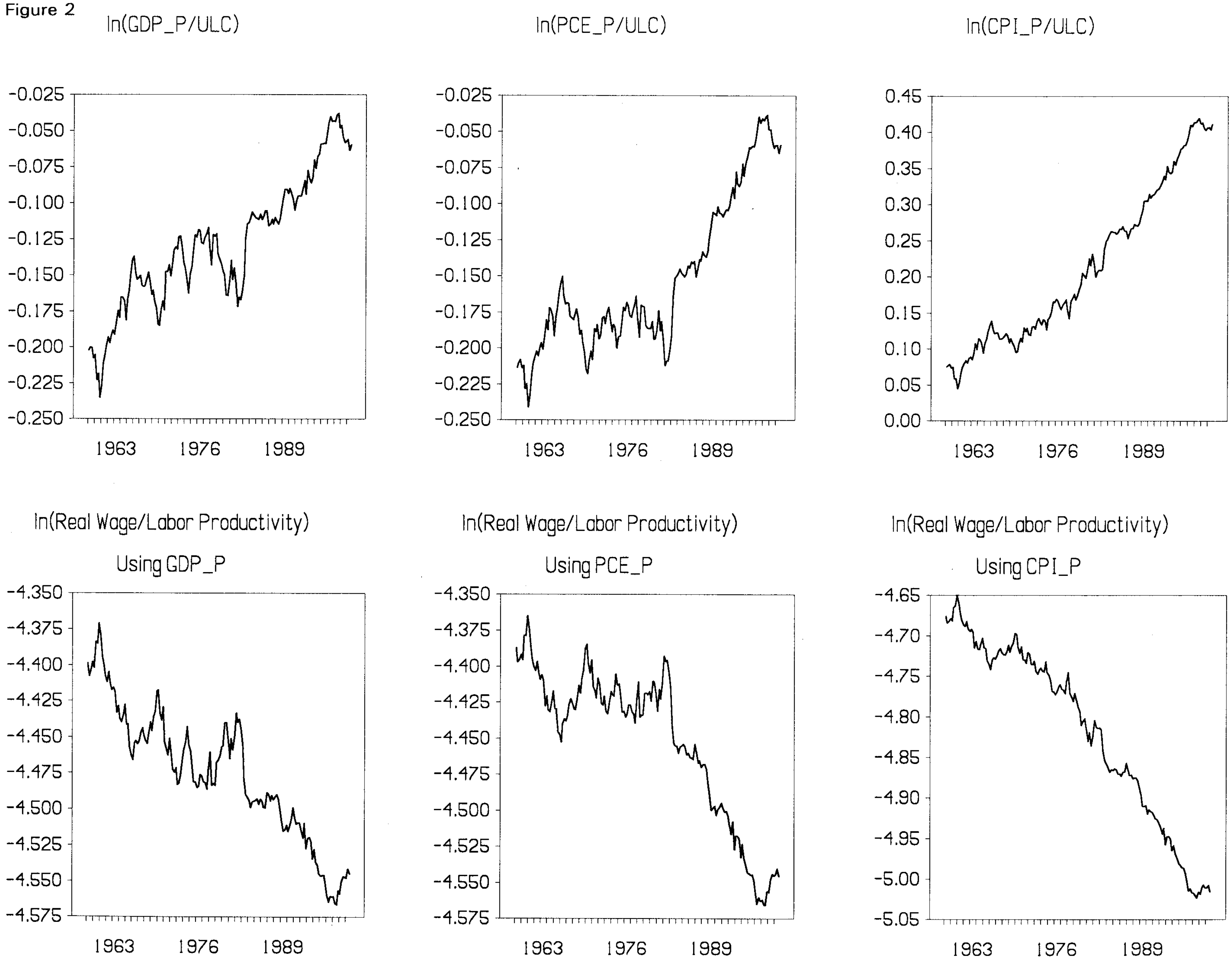

Growth in Unit Labor Costs
PCE Inflation
Growth in Compensation
CPI Inflation
Growth in Adjusted Monetary Base
GDP Inflation
Growth in Labor Productivity
1963
1975
1987

Figure 2


ln(GDP_P/ULC)
ln(PCE_P/ULC)
ln(CPI_P/ULC)
ln(Real Wage/Labor Productivity)
Using GDP_P
ln(Real Wage/Labor Productivity)
Using PCE_P
ln(Real Wage/Labor Productivity)
Using CPI_P
1963
1976
1989