JOURNAL OF APPLIED ECONOMETRICS, VOL. 10, 433–445 (1995)

CYCLICAL OUTPUT, CYCLICAL UNEMPLOYMENT, AND OKUN'S COEFFICIENT: A NEW APPROACH

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SUMMARY

Estimates of Okun's coefficient are obtained using new estimates of cyclical GNP and cyclical unemployment rates for the post-war USA. Empirical estimates of the coefficient are near −0·25, somewhat smaller in magnitude than other recent estimates obtained applying similar econometric techniques to different estimates of cyclical output and unemployment. Tests fail to reject the hypothesis of parameter stability across an hypothesized break between the third and fourth quarters of 1973, suggesting similar relationships between cyclical output and unemployment both before and after the supply shocks of the 1970s.

1. INTRODUCTION

Economists have long been aware of an inverse relationship between cyclical fluctuations in output and the level of or change in unemployment rates. This relationship is an obvious feature of the supply side of the macroeconomy: in the absence of shocks to productivity, as output rises in a cyclical recovery, some unemployed workers are hired to produce it. As output falls in recession, some workers are no longer needed and temporarily lose their jobs. However, the search for a quantifiable relationship suitable for policy analysis emerged comparatively recently in the long history of business cycle research with Okun (1962).

To fix ideas, let y_t and y_t^n represent, respectively, the logs of observed and potential GNP. Similarly, let U_t and U^n represent the observed and natural rates of unemployment.[1] We can write:

$$y_t^c \equiv y_t - y_t^n \tag{1}$$

$$U_t^c \equiv U_t - U^n \tag{2}$$

$$U_t^c = \alpha y_t^c \qquad \alpha < 0 \tag{3}$$

Here, y_t^c is the log output ratio; U_t^c is the cyclical unemployment rate, the difference between the observed unemployment rate and the natural rate. Thus, equation (3) is one version of Okun's law.

Here, we use a method based on the results of Perron (1989, 1990), Evans (1989a) and Weber (1993) to estimate the cyclical components of output and the unemployment rate. We then use these series to estimate Okun's coefficient and to test for its stability.

[1] No t subscript appears on U^n due to the assumption (discussed below) that, apart from a one-time break, U^n is constant throughout the sample period considered.

CCC 0883–7252/95/040433–13

Received July 1993
Revised February 1995

The paper is organized as follows. Section 2 reviews recent advances in measuring and interpreting Okun's coefficient. Section 3 presents the method used to estimate the cyclical components of output and the unemployment rate. The key question addressed is whether estimates of Okun's coefficient are sensitive to the method used to estimate cyclical output and unemployment. This is taken up in Section 4, which uses the estimates of cyclical output and unemployment from Section 2 to produce both static and dynamic estimates of α. Most of the estimates are rather small (less than 0·3 in magnitude), so that some doubt is cast on recent, larger estimates. Section 5 contains conclusions.

2. MEASUREMENT AND INTERPRETATION

Using an equation similar to equation (3), Okun (1962) estimated α to be near $-0{\cdot}32$.[2] While this estimate stood for over two decades as a useful empirical approximation, re-examinations of both the size and the proper interpretation of Okun's coefficient have appeared recently. The following extensions are of particular interest.

2.1. Theory

Gordon (1984) derives Okun's relationship from an identity relating the output ratio to the unemployment rate, labour force participation, productivity, average hours, and population. Thus, $-1{\cdot}0 < \alpha < 0$ implies that the fraction of output growth explained by factors other than the decline in unemployment is $1 + \alpha$.

Similarly, Prachowny (1993) shows that equation (3) is a special case of a production function which models output as a function of the unemployment rate, the capacity utilization rate, hours per worker, and the labour force. In this model, traditional versions of Okun's Law impose the condition that the cyclical components of capacity utilization, average hours, and the labour force do not affect the cyclical component of output. Since his empirical results reject this restriction, Prachowny suggests that earlier estimates of the coefficient were subject to an important specification bias.[3] He also finds that previous estimates of $1/\alpha$ ranging in magnitude from 2 (Gordon, 1984) to 3 (Okun, 1962) were too large. His estimates are between 0·6 and 0·7, which, he argues, is consistent with deriving Okun's relationship from an aggregate production function.

2.2. Dynamic Analysis

Gordon (1984) and Evans (1989a) emphasize a dynamic interpretation of the output–unemployment rate relationship and use autoregressive-distributed lag models to estimate the lagged effects of output growth on the unemployment rate. Both produce long-run estimates of Okun's coefficient exceeding 0·4 in magnitude; Gordon's estimate is nearly 0·5.

[2] Of course, Okun was careful not to rely on a single estimation method only. He also reports results from estimating the equation in first differences and in elasticity form. See Okun (1962).

[3] Prachowny's results may also bear negatively on the results reported below, which are based on a bivariate output–unemployment relationship, and do not account for the cyclical components of average hours, capacity utilization, or labour force participation. His results were brought to our attention after the original draft of this paper had been completed. In other work currently in progress, we combine Prachowny's methodology with that employed here to determine the extent to which his results depend on the method used to estimate cyclical output and unemployment rates and the extent to which they are affected by estimating dynamic models similar to equation (7). (Prachowny's model does not account for possible distributed lag or autoregressive relationships among his variables.)

2.3. Supply versus Demand Shocks

Blanchard and Quah (1989) use a long-run restriction on a reduced-form VAR in unemployment rates and output growth to obtain estimates of aggregate demand and supply innovations. They argue that Okun's coefficient is a mongrel coefficient, since the relationship between output growth and the unemployment rate depends on whether a shock to these series is a demand-side shock or a supply-side shock.

The usual explanation of Okun's coefficient interprets short-run shocks as driven purely by the demand side, so that output and unemployment move in opposite directions. However, a temporary shock to aggregate supply (productivity) could affect cyclical output without affecting the unemployment rate, or output and the unemployment rate could even move in the same direction. Thus, estimating Okun's coefficient involves an identification problem: shocks to aggregate demand and supply must be separated in order to identify a supply-side parameter such as Okun's coefficient. Blanchard and Quah find very different patterns of response of output and the unemployment rate to demand and supply shocks. For demand shocks, Blanchard and Quah find that Okun's coefficient is just under 0·5 in magnitude. They find no relationship between output growth and the unemployment rate following supply shocks.

2.4. Innovations in a VAR

Blanchard (1989) uses a five-variable VAR in GNP, unemployment, a price level, a wage, and money to test a short-run Keynesian macro model. He obtains a version of Okun's coefficient in the process of retrieving structural innovations in the VAR from the reduced form. Let e_{ut} and e_{yt} be innovations in the unemployment rate and output growth derived from a VAR. One version of Okun's coefficient is obtained from

$$e_{ut} = \alpha e_{yt} + u_t \tag{4}$$

Here, Okun's coefficient represents a relationship between VAR innovations rather than output growth and the unemployment rate themselves. In different parameterizations, Blanchard obtains estimates of Okun's coefficient between 0·18 and 0·42 in magnitude.

The findings of Gordon (1984), Evans (1989a), Blanchard and Quah (1989), and Blanchard (1989) have led Gordon (1993) to declare Okun's estimate of −0·32 obsolete. This is important, since the coefficient yields a forecast of the speed with which the unemployment rate moves over the business cycle. The recent larger estimates suggest that short-run output growth has a larger impact on unemployment rates than was once thought. For this reason, it is important to test the robustness of recent estimates to alternative specifications and alternative estimates of cyclical output and unemployment. The key question is: does output growth really exert such a large influence on the unemployment rate?

3. CYCLICAL OUTPUT AND UNEMPLOYMENT

One difficulty in estimating equation (3) lies in measuring U^n and y_t^n, which are unobserved. We follow Perron (1989) in measuring y_t^n and Evans (1989a), Perron (1990), and Weber (1993) in measuring U^n. Perron (1989) shows that deviations of log post-war US real GNP from a deterministic trend with a break in the trend coefficient during 1973 are stationary.[4] Deviations

[4] As Banerjee *et al.* (1992) and Zivot and Andrews (1992) have shown, the finding that real GNP is stationary around a break in the trend in 1973 depends sensitively on imposing the break as an exogenous event rather than permitting the data to determine the break date endogenously.

of the post-war US unemployment rate from a mean level which changes between the third and fourth quarters of 1973 are also stationary.[5] Thus, we assume that log potential output is a broken linear trend and that the natural rate of unemployment is the unconditional mean rate after accounting for a single change in mean in 1973. The deviations from these broken trends are used as estimates of the log output ratio and the cyclical rate of unemployment.[6] That is, we estimate y_t^n and U^n (y_t^c and U_t^c), respectively, as the fitted values (residuals) from the regressions

$$y_t = \beta_0 + \beta_1 t + \beta_2 D_t t + \varepsilon_{1t} \tag{5}$$

$$U_t = \gamma_0 + \gamma_1 D_t + \varepsilon_{2t} \tag{6}$$

where t is a time trend, D_t is a dummy variable equal to 0 up to and including the third quarter of 1973 and equal to one afterward, and ε_{1t} and ε_{2t} are stationary random disturbance terms.

Equations (5) and (6) were estimated using quarterly data for US real GNP and the civilian unemployment rate for 1948:1 to 1988:4. The parameter estimates imply an average annual rate of growth of 3·47% over 1948:1 to 1973:3 and 2·44% over 1973:4 to 1988:4 and a natural rate of unemployment of approximately 4·8% and 7·8% for the same periods. Estimates of the output gap and the cyclical unemployment rate are shown in Figures 1 and 2; the vertical lines

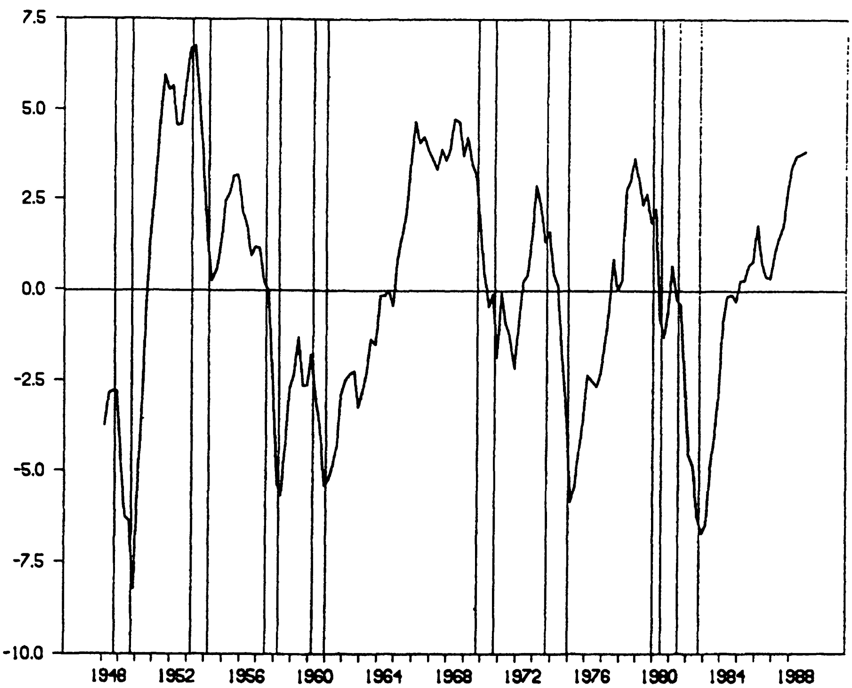

Figure 1. Cyclical real GNP

[5] See Weber (1993). Evans (1989a) and Perron (1990) produce similar results except that they permit the mean unemployment rate to change following the fourth, rather than the third quarter of 1973.

[6] Okun (1962) used similar methods to measure the cyclical output and the cyclical unemployment rate, but assumed no change in either the growth rate of potential output or the natural rate of unemployment.

Table I. Unit root tests

Equation: $x_t = \rho x_{t-1} + a + bt + cD_t + dD_t t + \sum_{i=1}^{j} \Delta x_{t-i} + \varepsilon_t$

Series	Specification	DF
y_t	$c=d=0$	$-1{\cdot}419$
y_t	Unrestricted	$-3{\cdot}817^{a}$
U_t	$b=c=d=0$	$-2{\cdot}057$
U_t	$b=d=0$	$-4{\cdot}013^{b}$

[a,b] Significance at the 0·10 and 0·05 levels, respectively. The unit root t-statistics for y_t and U_t are very nearly significant at the 0·05 and 0·01 levels, respectively. Data are monthly for the unemployment rate, quarterly for log real GNP.

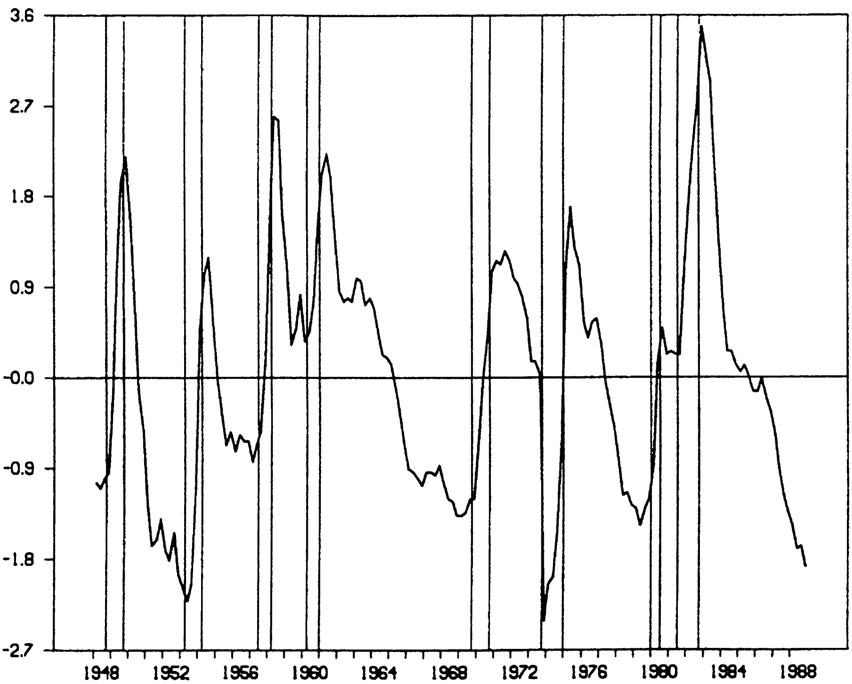

Figure 2. Cyclical unemployment rate

show NBER business cycle peaks and troughs. The peaks and troughs of both series coincide well with NBER business cycle peaks and troughs.

Unit root test results are reported in Table I. These confirm results of Perron (1989, 1990), Evans (1989a), and Weber (1993). In levels, y_t and U_t both contain a unit root, but imposing a break in the trend growth rate of y_t in October, 1973 renders detrended output stationary, while imposing a break in the mean unemployment rate in October, 1973 renders the unemployment rate stationary.

4. RESULTS

4.1. Estimation Methods

We use four different methods to estimate Okun's coefficient. The first is static OLS applied to equation (3). Second, a long-run version of the coefficient can be estimated using a method of Gordon (1984). First, an autoregressive-distributed lag model is estimated for the unemployment rate:

$$U_t^c = \sum_{i=1}^{k} \delta_{1i} U_{t-i}^c + \sum_{i=1}^{k} \delta_{2i} y_{t-i}^c + \varepsilon_{3t} \tag{7}$$

Let d_{1i} and d_{2i} be estimates of δ_{1i} and δ_{2i}. Then a^{LR}, an estimate of the impact of a change in y_t^c on U_t^c in the long run, is

$$a^{\mathrm{LR}} = \frac{\sum_{i=1}^{k} d_{2i}}{1 - \sum_{i=1}^{k} d_{1i}}$$

Equation (7) is estimated using quarterly data and two values for k: $k=2$ and $k=4$. The results are robust to the choice of k.

The third method is adapted from Blanchard (1989). This method estimates a version of α relating innovations in the cyclical unemployment rate to innovations in cyclical output. A two-step procedure is used: First, estimate a bivariate VAR, including equation (7) and the analogous AR-DL equation for y_t^c:

$$y_t^c = \sum_{i=1}^{k} \delta_{3i} U_{t-i}^c + \sum_{i=1}^{k} \delta_{4i} y_{t-i}^c + \varepsilon_{4t} \tag{8}$$

Using e_{3t} and e_{4t}, which are estimates of ε_{3t} and ε_{4t}, Okun's coefficient is estimated in the process of retrieving structural innovations from their reduced-form counterparts by estimating[7]

$$e_{3t} = ae_{4t} + u_t \tag{9}$$

where u_t is the portion of the cyclical unemployment rate innovation which is orthogonal to the cyclical GNP innovation.

Finally, Okun's Law suggests that after detrending, y_t and U_t should be cointegrated, so that the cointegrating regression yields a 'super-consistent' estimator of α (Engle and Granger, 1987).[8] To see this, solve equations (5) and (6) for y_t^c and U_t^c, and substitute into equation (3). This yields, after rearranging

$$U_t = \eta_0 + \eta_1 D_t + \eta_2 t + \eta_3 D_t t + \alpha y_t + \varepsilon_{5t} \tag{10}$$

[7] In fact, there are several differences between the method used by Blanchard and that employed here: First, Blanchard renders y_t stationary by removing a deterministic trend from its first difference; he renders U_t stationary by deterministic detrending. Second, Blanchard uses a five-variable VAR, including a price level, a wage rate, and a money variable in addition to output and the unemployment rate. Third, Blanchard reports several estimates of Okun's coefficient in which supply innovations are permitted to affect aggregate demand in addition to the simpler model (analogous to that used here) in which aggregate supply innovations do not influence aggregate demand.

[8] This approach was suggested by an anonymous referee.

where $\eta_0 \equiv \gamma_0 - \alpha\beta_0$, $\eta_1 \equiv \gamma_1$, $\eta_2 \equiv -\alpha\beta_1$, and $\eta_3 \equiv -\alpha\beta_2$. Thus, a fourth method is: estimate the cointegrating regression; test the residuals for stationarity; if they are stationary, the OLS estimate of α is a long-run estimate of Okun's coefficient.[9]

It should be noted that these methods represent different conceptual approaches to the output–unemployment rate relationship, not different methods of estimating the same theoretical construct. Thus, static OLS assumes a strictly contemporaneous relationship, ignoring both short- and long-run dynamics. Equation (7) estimates a dynamic short-run relationship over a year or less. Blanchard's method measures the relationship between one-step-ahead forecast errors in a forecasting model. The cointegrating regression estimates a long-run equilibrium relationship, explicitly ignoring the transitional dynamics captured in equation (7). Any comparison of the empirical results below must bear in mind that they are not different estimates of the same parameter but estimates of the output–unemployment relationship at three different time horizons, contemporaneous, short-run, and long-run and of the contemporaneous relationship between output and unemployment rate innovations.

Finally, since several economic 'constants', including the natural rate of unemployment and the long-run rate of growth of GNP apparently changed in 1973,[10] we test whether Okun's coefficient changed then as well. For the OLS estimates, this hypothesis is tested by adding D_t to equation (3) as a slope dummy variable:

$$U_t^c = \alpha_1 y_t^c + \alpha_2 D_t y_t^c + \varepsilon_{1t} \tag{3'}$$

Okun's coefficient is α_1 before the break and $\alpha_1 + \alpha_2$ afterwards. Parameter constancy implies $\alpha_2 = 0$. The same method is used for estimates from Blanchard's method or the cointegrating regression.

For the dynamic OLS estimates, the same slope dummy was added to each variable on the right-hand side of equation (7), yielding:

$$U_t^c = \sum_{i=1}^{k} \delta_{1i} U_{t-i}^c + \sum_{i=1}^{k} \delta_{2i} y_{t-i}^c + \sum_{i=1}^{k} \delta_{11i} D_t U_{t-i}^c + \sum_{i=1}^{k} \delta_{21i} D_t y_{t-i}^c + \varepsilon_{3t} \tag{7'}$$

From equation (7′), the long-run estimates of Okun's coefficient are

$$a^{\mathrm{LR}} = \frac{\sum_{i=1}^{k} d_{2i}}{1 - \sum_{i=1}^{k} d_{1i}}$$

[9] The fact that U_t and y_t should be cointegrated raises an important technical question regarding equation (7) and the VAR consisting of equations (7) and (8). Specifically, Engle and Granger (1987) show that if two variables are separately difference stationary but cointegrated, then a VAR in first differences is misspecified, since it should contain the error-correction term, the series of errors from the cointegrating regression. (If three or more difference stationary variables are cointegrated, there may be more than one cointegrating regression, more than one long-run equilibrium relationship, and more than one error-correction term.) However, the question of including the errors from the cointegrating regression (10) in equations (7) and (8) does not arise here, since the unemployment rate and log output are rendered stationary by deterministic detrending (equations (5) and (6)), not by first differencing. That is, equations (7) and (8) are already an error-correction mechanism: for parameter values consistent with asymptotic stability, they determine how the cyclical (disequilibrium) levels of U_t and y_t converge asymptotically to zero. Thus, including the error-correction terms from equation (9) would be redundant.

[10] See e.g. Perron (1989, 1990), Evans (1989a), and Weber (1993).

and

$$a^{\mathrm{LR}} = \frac{\sum_{i=1}^{k} (d_{2i} + d_{21i})}{1 - \sum_{i=1}^{k} (d_{1i} + d_{11i})}$$

before and after the break, respectively. Since a^{LR} is a nonlinear combination of the coefficients in equation (7′), the standard errors and covariances used to test the hypothesis of a break in a^{LR} are estimated using an approximation suggested by Greene (1993, p. 75).[11]

4.2. Empirical Results

Table II, reports eighteen estimates of Okun's coefficient: six estimation methods were used—static OLS, the cointegrating regression, dynamic OLS with two and four lags, and Blanchard's method using two and four lags in the VAR. Each method was applied to three different sample periods, the full sample, 1948:1 to 1988:4 and two shorter samples, 1948:1 to 1973:3 and 1973:4 to 1988:4.

From Table II, if cyclical unemployment and output are estimated in the manner described above, Okun's estimate of the coefficient is too large, not too small: thirteen out of eighteen estimates are less than −0·3 in magnitude; only two exceed −0·4.

The static OLS estimate of −0·314 is very close to Okun's own estimate of −0·32,[12] but the OLS results reject parameter constancy: the t-statistic is 3·580, and the difference in parameter estimates around the split is substantial (−0·277 to −0·402), suggesting a large, statistically significant increase in magnitude.[13]

For the cointegrating regression, we obtained critical values for the ADF statistic for the null hypothesis of non-cointegration by simulation.[14] The ADF statistic for the null of non-cointegration around a changing mean and trend is −5·10, which is significant at the 0·01 level. The errors from the cointegrating regression are shown in Figure 3.

The estimates of α from the cointegrating regression are similar to those from the static regression. Parameter constancy is again rejected: the t-statistic is 5·628, and the difference in parameter estimates is substantial: −0·293 to −0·465.

However, the low Durbin–Watson statistics for both regressions suggest first-order serial correlation in the residuals, while the small Goldfeld–Quandt statistics imply heteroscedasticity with a greater residual variance in the earlier sub-sample than in the latter. These findings cast doubt on the standard errors and the stability test results, and call for an alternative specification. Following Sargan (1964) and Hendry and Mizon (1978), we take autocorrelation in the errors as evidence of misspecified dynamics, so that the estimates may be subject to specification bias. For this reason, we consider next the dynamic equations (7) and (7′).

The estimates, labelled dynamic OLS in the table, range from −0·234 for the early sample period using $k=2$ lags to −0·333 for the later sample period also using $k=2$ lags. These

[11] A longer version of this paper, which is available on request, contains a detailed description of the method used.
[12] This is all the more interesting since Okun in 1962 took the natural rate of unemployment to be somewhat lower than that used here for the period 1948:1 to 1973:3 (4% rather than 4·8%).
[13] In a somewhat different specification, Prachowny (1993) finds parameter instability around an hypothesized break at the end of 1974.
[14] The Monte Carlo simulation is described in detail in a longer version of this paper. The lower tail critical values were: 0·01: −5·054; 0·05: −4·505; 0·10: −4·256.

Table II. Estimates of Okun's coefficient. Quarterly data: 1948:1 to 1988:4

Estimation method and period[a]	a[b]	R^2[c]	Chow[d]	G–Q[e]	DW/LM[f]
Static OLS					
Full sample	−0·314	0·689		0·293	0·266
	(0·017)[f]			(1·000)	
1948:1 to 1973:3	−0·277	0·712			0·281
	(0·019)				
1973:4 to 1988:4	−0·402	0·712	3·580		0·281
	(0·029)		(0·000)		
Cointegrating regression[g]					
Full sample	−0·340	0·878		0·286	0·279
	(0·015)			(1·000)	
1948:1 to 1973:3	−0·293	0·899			0·345
	(0·016)				
1973:4 to 1988:4	−0·465	0·899	5·628		0·345
	(0·026)		(0·000)		
Dynamic OLS, $k = 2$[h]					
Full sample	−0·260	0·914		0·961	2·489
	(0·038)			(0·553)	(0·647)
1948:1 to 1973:3	−0·234	0·917			2·570
	(0·039)				(0·632)
1973:4 to 1988:4	−0·333	0·917	1·166		2·570
	(0·075)		(0·245)		(0·632)
Dynamic OLS, $k = 4$					
Full sample	−0·260	0·916		0·992	2·326
	(0·039)			(0·502)	(0·676)
1948:1 to 1973:3	−0·252	0·921			2·700
	(0·035)				(0·609)
1973:4 to 1988:4	−0·257	0·921	0·014		2·700
	(0·114)		(0·989)		(0·609)
Blanchard, $k = 2$					
Full sample	−0·224	0·350		0·996	2·060
	(0·024)			(0·498)	
1948:1 to 1973:3	−0·195	0·360			2·049
	(0·030)				
1973:4 to 1988:4	−0·275	0·360	1·613		2·049
	(0·040)		(0·109)		
Blanchard, $k = 4$					
Full sample	−0·224	0·347		0·896	1·983
	(0·024)			(0·664)	
1948:1 to 1973:3	−0·190	0·361			1·981
	(0·030)				
1973:4 to 1988:4	−0·284	0·361	1·878		1·981
	(0·040)		(0·062)		

[a] Sub-sample results (1948:1 to 1973:3 and 1973:4 to 1988:4) are from single equations with slope dummy variables added for the period 1973:4 to 1988:4.
[b] Standard errors in parentheses. For a^{LR} (dynamic OLS), standard errors are approximations calculated using first-order Taylor expansions as in Greene (1993, p. 75).
[c] For dynamic OLS, R^2 is reported for the original equation for the unemployment rate.
[d] Chow is the t-statistic for parameter constancy against the alternative of a break at 1973:3; p-values in parentheses.
[e] G–Q denotes the Goldfeld–Quandt test for heteroscedasticity; p-values in parentheses. Sixteen observations centred at 1973:3/1973:4 were omitted.
[f] DW and LM are the Durbin–Watson statistic for first-order serial correlation and the Breusch–Godfrey LM statistic for serial correlation up to order 4 (asymptotically $\chi^2(4)$ under the null hypothesis of no serial correlation), respectively. DW is reported for static OLS, Blanchard, and cointegrating regression; LM is reported for dynamic OLS. For LM, p-values are in parentheses.
[g] The augmented Dickey–Fuller statistic for cointegration of log output and the unemployment rate around a broken trend and a break in mean is −5·10, which is significant at the 0·01 level. The lower tail critical values, obtained by simulation, were 0·01: −5·054; 0·05: −4·505; 0·10: −4·256.
[h] k denotes lag length for autoregressive-distributed lag equations.

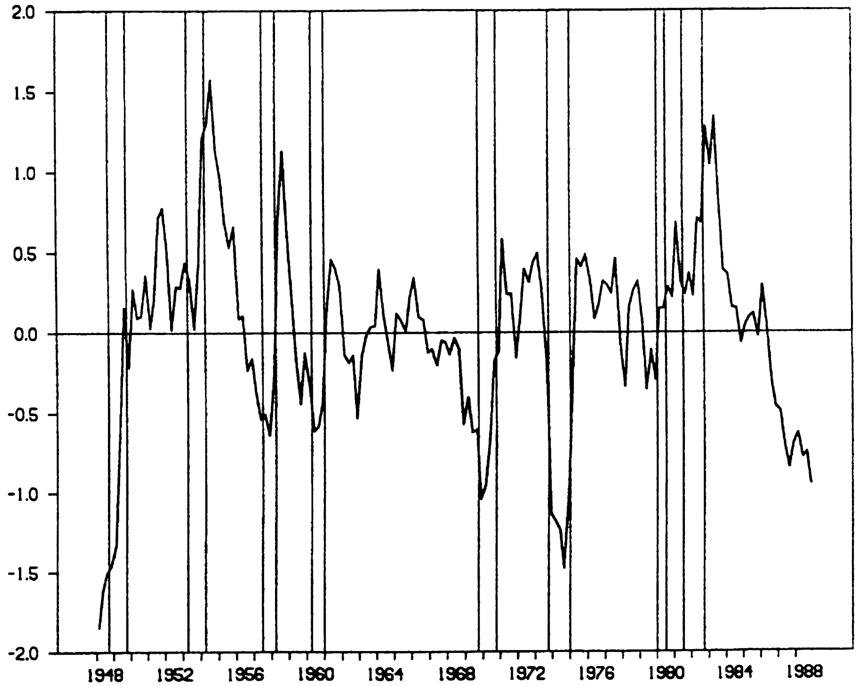

Figure 3. Error from cointegrating regression

estimates are universally smaller than their counterparts from the static and cointegrating regressions: for both lag lengths, the full sample estimate of the long-run Okun coefficient is $-0{\cdot}26$. Also, the t-statistics for parameter stability[15] now strongly fail to reject the null hypothesis.[16] Serial correlation of the residuals appears not to be a problem in the dynamic model. For both two and four lags, the LM statistics for autocorrelation of order up to four all fail to reject the null hypothesis of no serial correlation.

Finally, we consider results obtained using Blanchard's method. The Durbin–Watson test for serial correlation is valid here since no lags of the dependent variable appear on the right-hand side. The D–W statistics fail strongly to reject the null hypothesis of no first-order serial correlation in the residuals.[17] Again, the t-statistic for the model with two lags in the initial VAR[18] fails to reject the null hypothesis of parameter constancy, although here the rejection is not quite as strong. Consistent with the results above and those of Blanchard (1989), the estimate of α is rather small: $-0{\cdot}224$ for the full sample.[19] For the estimate

[15] The methods for calculating approximate standard errors for the estimates of α^{LR} and the test statistics for stability of α^{LR} are discussed in a longer version of this paper.

[16] This confirms results reported by Evans (1989a,b).

[17] This is not surprising since the dynamics of the unemployment–output relationship are already contained in equations (7) and (8), which are estimated prior to estimating equation (9).

[18] The tests here are direct tests for the stability of α in equation (9) rather than tests for the stability of the lag coefficients in the initial VAR.

[19] Blanchard's estimate of $-0{\cdot}18$ requires the identifying assumption that supply shocks do not influence aggregate demand. When supply shocks are allowed to influence aggregate demand the estimate rises. Using somewhat more sophisticated methods, Blanchard reports an estimate of Okun's coefficient of $-0{\cdot}42$, which is in the neighbourhood of other recent estimates (see above).

obtained from the four-lag VAR, the estimates are quantitatively similar: the full sample estimate is $-0{\cdot}224$, with an estimate of $-0{\cdot}190$ for the early sub-sample and $-0{\cdot}284$ for the later one. However, in this case, the t-statistic for parameter constancy is larger (1·878), with a marginal significance level less than 0·07. Thus, the estimate derived from the four-lag VAR does yield some limited evidence of a significant increase in the coefficient in 1973.

4.3. Discussion

Combined with generally larger estimates of Okun's coefficient obtained elsewhere using different estimates of cyclical output and the cyclical unemployment rate, the results above suggest that estimates of the coefficient are sensitive to the method used to estimate cyclical output and unemployment. This sensitivity is found even when similar econometric procedures are used.

Also, the stability of the dynamic version of the coefficient sheds light on the empirical importance of Blanchard and Quah's (1989) argument that Okun's coefficient is a mongrel coefficient. While the theoretical validity of Blanchard and Quah's argument is not questioned here, their criticism appears inconsistent with the results reported above. If fluctuations in output and unemployment rates prior to 1973 were driven largely by demand shocks, while those of the mid- to late 1970s were due more to supply shocks, then Blanchard and Quah's criticism predicts that Okun's coefficient should be larger for the earlier period than for the latter. This is inconsistent with the results in Table II. For this reason, we consider two further tests of the empirical importance of their criticism of Okun's coefficient.

First, since the focus of the criticism is on the source of the shock to output (i.e. on whether the shock affects aggregate demand or aggregate supply) we consider tests for stability of α^{LR} when only the distributed lag coefficients on output are permitted to change. (That is, with the autoregressive coefficients on the unemployment rate assumed constant.) If the output shocks of the 1970s were different in nature from those which preceded them while the relationship between current and past rates of unemployment remained unchanged, then in the autoregression for the unemployment rate only the distributed lag coefficients on output should change. Results of tests for stability of α^{LR} with the AR coefficients held constant are reported in panel A of Table III.

Second, on the assumption that supply shocks dominate only the period between 1973 and 1979 (the period between the first oil shock and the October, 1979 change in Federal Reserve Board policy), we consider tests for parameter stability over the period 1973:4 to 1979:3 with the parameters assumed to have the same value over the earlier and later periods. Results of tests for stability over this middle period are reported in both panels of Table III.

All six t-statistics are still statistically insignificant, but in three cases the change in the coefficient is at least in the right direction. The coefficient is smaller over the 1973:4 to 1979:3 sub-period. Assuming that this period was dominated by supply shocks, this is consistent with Blanchard and Quah's criticism.[20]

[20] While the statistical insignificance of this result is a concern, it may simply suggest that demand shocks were not entirely absent over the sub-period 1973:4 to 1979:3, while supply shocks were not entirely absent over the earlier and later periods. That is, it is unlikely that we have managed to isolate periods dominated solely by either demand or supply shocks.

Table III. Additional long-run estimates of Okun's coefficient. Quarterly data: 1948:1 to 1988:4

Estimation method and period	a	R^2	Chow	LM
A. Coefficients on lagged output only permitted to change				
Dynamic OLS, $k = 2$				
Full sample	−0·260	0·914		2·489
	(0·038)			(0·647)
1948:1 to 1973:3	−0·239	0·915		3·409
	(0·041)			(0·492)
1973:4 to 1988:4	−0·327	0·915	1·183	3·409
	(0·065)		(0·239)	(0·492)
1948:1 to 1973:3/	−0·266	0·915		2·521
1979:4 to 1988:4	(0·038)			(0·641)
1973:4 to 1979:3	−0·213	0·915	−0·482	2·521
	(0·106)		(0·630)	(0·641)
Dynamic OLS, $k = 4$				
Full sample	−0·260	0·916		2·326
	(0·039)			(0·676)
1948:1 to 1973:3	−0·251	0·917		2·411
	(0·043)			(0·661)
1973:4 to 1988:4	−0·295	0·917	0·534	2·411
	(0·072)		(0·594)	(0·661)
1948:1 to 1973:3/	−0·272	0·918		2·170
1979:4 to 1988:4	(0·039)			(0·704)
1973:4 to 1979:3	−0·169	0·918	−0·884	2·170
	(0·112)		(0·704)	(0·704)
B. All lag coefficients permitted to change				
Dynamic OLS, $k = 2$				
Full sample	−0·260	0·914		2·489
	(0·038)			(0·647)
1948:1 to 1973:3/	−0·255	0·917		1·572
1979:4 to 1988:4	(0·041)			(0·814)
1973:4 to 1979:3	−0·262	0·917	0·057	1·572
	(0·107)		(0·955)	(0·814)
Dynamic OLS, $k = 4$				
Full sample	−0·260	0·916		2·326
	(0·038)			(0·676)
1948:1 to 1973:3/	−0·274	0·922		1·954
1979:4 to 1988:4	(0·040)			(0·744)
1973:4 to 1979:3	−0·178	0·922	−0·400	1·954
	(0·131)		(0·690)	(0·744)

See notes to Table II.

5. CONCLUSIONS

Drawing on results of Perron (1989, 1990), Evans (1989a), and Weber (1993), we have estimated the cyclical portions of post-war US real GNP and unemployment rates. These series were used to estimate Okun's coefficient. While the static estimates support Okun's (1962) original estimate near −0·32 and indicate an increase in the coefficient in 1973, four different dynamic specifications all suggest a smaller value for the coefficient. Also, the dynamic results generally support parameter constancy across an hypothesized break in 1973. This confirms similar results of Evans (1989a,b).

The methods used here to estimate the cyclical components of GNP and the unemployment rate are not the only ones possible. It would be interesting to know whether estimates obtained using other methods yield results similar to those reported here or more like the results of previous studies cited above.

Also, these results are difficult to reconcile with Blanchard and Quah's (1989) criticism of Okun's coefficient as a mongrel coefficient. Further research might help determine the extent to which the dynamic relationship between cyclical output and cyclical unemployment rates depends on the initial source (demand side or supply side) of a particular cyclical fluctuation.

ACKNOWLEDGEMENTS

I am grateful to two anonymous referees for helpful comments on an earlier draft of this paper. Naturally, all errors are my own responsibility.

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