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Alleged instability of the Okun's law relationship in Australia: an empirical analysis

MARTIN WATTS and WILLIAM MITCHELL

Department of Economics, University of Newcastle, New South Wales 2308, Australia

The determinants of the long run equilibrium unemployment rate are modelled through a cointegrating regression. The residual term, which represents the error correction mechanism, is used in the testing of a general dynamic model to obtain a simplified representation of the data generation process over the period 1968(2)–1989(2). The long period equilibrium unemployment rate, which is $I(1)$, is shown to be related to the rate of capacity utilization, a structural change variable and the rate of capacity growth. Steady state rates of unemployment are calculated which are based on different assumptions about the magnitudes of the independent variables. In contrast to the other two variables, changes in the rate of capacity utilization are shown to have only modest effects on the steady state unemployment rate. The Okun coefficient is inversely related to the steady state unemployment rate, which accords with intuition. The statistically significant decline in the trend growth rate of non-farm GDP in the 1970s explains, in part, the apparent instability of the Okun's law relationship which is revealed in recent studies. These studies can be criticised for their econometric methodologies.

I. INTRODUCTION

Recent Okun's law studies in Western economies have pointed to the instability of the relationship in the late 1960s and early 1970s. For example, Thurow (1983, p. 9) and Okun (1980, p. 168) make this claim with respect to the US and Davenport (1982) argues that the breakdown of the relationship occured in Canada in 1967. Likewise, recent Australian studies by the Bureau of Labour Market Research (BLMR) (1987) and Nguyen and Siriwardana (1988) argue that structural shifts have influenced the unemployment–output gap nexus, where the output gap is defined as the inverse of the level of capacity utilization minus unity multiplied by 100.

In this study, we argue that Okun's law, viewed as a long run time series relationship, has been mis-specified from the outset, and that its initial stability was a consequence of the overall economic stability which existed at the time. Further, the previous studies which have explored relations between variables in level form are open to the criticism that the regressions between non-stationary economic variables may be spurious. In contrast, we employ the two-stage Engle–Granger cointegration procedure to overcome the spurious regression problem.

Further, Okun's original work and subsequent studies by others have all calculated the Okun coefficient on the basis of an assumed 'natural rate of unemployment'. Our work is different because, through the construction of a potential output series, we are able to estimate the Okun coefficient and the steady-state unemployment rate jointly, thus overcoming some of the arbitrariness of the other techniques.

We model the determinants of the long-run equilibrium unemployment rate via a cointegrating regression (CR). The residual from the CR is the measure of disequilibrium between the variables. Used as a regressor in a dynamic model, the residual term represents the error-correction (EC) mechanism. We also test down from a general dynamic model to achieve an adequate simplified representation of the data generation process over the period 1968(2)–1989(2), and confirm the robustness of the EC specification. Thus, short-run changes in the aggregate unemployment rate are, in part, considered to be adjustments to a steady-state level

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following changes to other influential variables such as economic activity. We show that our equilibrium unemployment rate estimates, derived from the CR, are compatible with the steady-state solution to the dynamic tested-down specification, which increases our confidence in the EC representation.

The long period equilibrium level of the unemployment rate is shown to be related to the rate of capacity utilization, a structural change variable, and the rate of capacity growth (all variables are described in the Appendix).[1]

II. LITERATURE

Okun (1962) hypothesized a relationship between real output and unemployment, although there has been debate over the direction of this relation. Of interest is the output level which accompanies the potential unemployment rate, namely the potential output level. Okun inverted this focus, however, by calculating the output gain that is necessary to achieve a target low unemployment rate. He viewed the unemployment rate as a summary of the various ways that slackness affected output.

He argued that the deviation of the unemployment rate about its steady state level was related to the level of excess demand in the product market. As output approached its potential value, employment rose to produce the extra output, and *ceteris paribus*, unemployment fell. The excess demand in the product market can be proxied by the gap between output (A) and potential output (P) expressed as a percentage of A.

With U^* denoting the steady-state unemployment rate, Okun's relation can be written as:

$$U-U^*=q[(P-A)/A]\times 100=q.YGAP \quad q>0 \qquad (1)$$

where q is the inverse of the Okun coefficient. The interpretation of Equation 1 is that if output is at potential ($A=P$), then $U=U^*$, whereas if the output gap is 1% of its actual level, say, then the unemployment rate would be q percentage points higher, *ceteris paribus*. The problem associated with this approach is that P is not observable. Okun generated a number of synthetic series and used the one which maximized the R^2.

If the parameters q, U^* are constant, then the relationship is stable. A constant rate of capacity utilization A/P, or, in other words, the equality of the actual and warranted (potential) rates of output growth implies a constant unemployment rate, and hence the equality of the actual and natural rates of growth.

Okun's law would only be expected to hold in this form for modest movements about the trend rate of potential output growth. In the 1970s there was a discrete change in the trend rate of growth of seasonally unadjusted, non-farm GDP (GNFP) and, by construction, the growth of capacity.[2] Slower capacity growth can be expected to lead to a rise in the rate of unemployment, in the long run, even if capacity utilization remains constant.

This would suggest that the unemployment rate is influenced by long-run cyclical factors, in addition to short-run variations in capacity utilization (Davenport, 1982).[3] Further, the process of structural change which affects the industrial composition of employment and its composition by sex and part-time and full-time status, would also be expected to influence the unemployment/capacity utilization nexus in the long run (Lilien, 1982; Abraham and Katz, 1984). For example, the increased share of female and part-time employment has repercussions for the Okun's Law relationship because female labour force participation differs from that of males (see below).

The Australian studies, such as BLMR (1987) and Nguyen and Siriwardana (1988), share a common characteristic in that their econometric methodologies are unspecified. No attempt is made to investigate the order of integration of the time series data and no justification is provided for the chosen specification (Watts and Mitchell, 1990).

These papers are representative of the approach usually adopted in Okun's law studies (see also Davenport (1982) and Gordon (1984)) in which the calculation of the Okun coefficient is often given priority over the specification of a long-run unemployment relationship with good econometric properties. Most importantly, the problems of spurious regression in a time series regression involving variables in levels and time trends are ignored.

III. UNIT ROOT TESTING AND COINTEGRATION METHODOLOGY

Important developments have occurred in the area of dynamic modelling over the last decade, especially in relation to the existence of cointegrated vectors and their relevance to EC models (Granger, 1981; Engle and Granger, 1987; Johansen 1988; Johansen and Juselius, 1990). Related developments have also been made in the area of unit root testing (Dickey and Fuller, 1979, 1981; Evans and Savin, 1981, 1984; Nelson and Plosser, 1982; Phillips, 1987; Phillips and Perron, 1986; Perron, 1988, 1989).

[1]To eliminate any cyclical supply effects on the unemployment rate due to added worker and/or discouraged worker effects, we use a trend labour force measure as the denominator in the calculation of the unemployment rate.

[2]The estimation of a simple trend through quarterly GNFP observations over the period 1966(3)–1973(4) yields GNFP $=25159.3+457.7\times t$, whereas over the remaining period 1974(1)–1989(2), GNFP $=25233.4+360.3\times t$. A Chow test of the stability of the regression coefficients over the sample period fails, $F(2,88)=5.54$, compared with a 5% critical value of 3.10.

[3]Robinson (1978, pp. 16–18) argues that the general notion of a distinction between changes in utilization and changes in productive capacity is indispensible for the analysis of industrial activity.

Table 1. *Sample autocorrelations (ACF) of the natural logs and first differences, 1966(4)–89(2)*

Series	1	2	3	4	5	6	7	8	9	10
LUR	0.97	0.94	0.91	0.89	0.85	0.81	0.79	0.77	0.74	0.71
ΔLUR	0.12	−0.12	0.09	0.23	−0.28	−0.36	−0.02	0.12	−0.09	−0.10
LCU	0.74	0.68	0.58	0.56	0.41	0.36	0.27	0.28	0.19	0.21
ΔLCU	−0.42	0.07	−0.14	0.30	−0.21	0.08	−0.21	0.23	−0.24	0.15
LSIG	0.54	0.49	0.52	0.58	0.47	0.46	0.51	0.43	0.39	0.42
$\Delta LSIG$	−0.45	−0.09	−0.04	0.17	−0.10	−0.07	0.16	−0.05	−0.06	−0.02
LGD	0.96	0.92	0.88	0.84	0.79	0.76	0.72	0.68	0.65	0.61
ΔLGD	0.94	0.89	0.83	0.78	0.72	0.67	0.61	0.56	0.51	0.45

In this section we establish initially the properties of the individual time series prior to testing for cointegration. Series which are integrated of a different order cannot be cointegrated.

Testing for unit roots in time series variables

Four seasonally-adjusted, quarterly time series are used:

(1) *LUR*, the log of the aggregate unemployment rate;
(2) *LGD*, the log of the potential output level based on linked peaks of non–farm, real domestic product;
(3) *LCU*, the log of the corresponding capacity utilization series;
(4) *LSIG*, the log of the standard deviation of the employment growth across the ASIC industrial structure.

Table 1 reveals some variation in the sample autocorrelations of the time series in level form. *LUR* and *LGD* are similar, starting at around 0.97 at lag one and slowly decaying as the lag increases. The ACF of a random walk displays a similar pattern to this (Nelson and Plosser, 1982, p. 147). *LSIG* and *LCU* have smaller autocorrelation coefficients at lag one, but they also decay smoothly and slowly.

Table 1 also reports the ACFs for the variables in first differences. *LGD* apart, the first lag is significant for all the variables, but at higher lags the functions drop off rapidly. This behaviour is consistent with stationarity. The first difference of *LGD* decays slowly with a high autocorrelation at lag one, which indicates that second-differencing is required.[4] The damping pattern of the ACFs for ΔLUR, ΔLCU and $\Delta LSIG$ suggest that an autoregressive component of order greater than one is present in these series. Only ΔLUR has any significant spikes indicating that the moving average component could be of order greater than one.

We now turn to more formal testing. There are various autoregressive representations which can be used as the basis for unit root testing. For example, Perron (1988) defines three regression equations, which indicate an ordering of relevant hypotheses.

$$y_t = \hat{\alpha} y_{t-1} + \hat{u}_t \quad (2)$$

$$y_t = \mu^* + \alpha^* y_{t-1} + u_t^* \quad (3)$$

$$y_t = \tilde{\mu} + \tilde{\beta}(t - T/2) + \tilde{\alpha} y_{t-1} + \tilde{u}_t \quad (4)$$

Equation 2 is stationary if $|\alpha| < 1$, whereas if $\hat{\alpha} = 1$, the process has a unit root and is non-stationary (Dickey and Fuller, 1979, p. 427, equation 1.1). Equation 3 allows for a fixed drift, μ^* (Dickey and Fuller, 1979, p. 428, equation 2.1). Equation 4 provides the framework for testing:

(1) the null of a driftless random walk $(\tilde{\mu}, \tilde{\beta}, \tilde{\alpha}) = (0, 0, 1)$ against the null being untrue;
(2) the null $(\tilde{\mu}, \tilde{\beta}, \tilde{\alpha}) = (\tilde{\mu}, 0, 1)$ (Dickey and Fuller, 1981, p. 1057, equation 1.3).

Nelson and Plosser (1982) compare trend-stationary (TS) to difference-stationary (DS) processes. The linear model which nests both hypotheses (as alternatives) is

$$y_t = \gamma + \phi t + u_t/(1 - \alpha L) \quad (5)$$

where L is the lag operator. Rearrangement of Equation 5 gives Equation 4 where $\mu = [\gamma(1-\alpha) + \alpha\phi)]$ and $\beta = \phi(1-\alpha)$. Thus under the null of a unit root, i.e. $\tilde{\alpha} = 1$ in Equation 4, the implied value of $\tilde{\beta}$ is zero, with the unit root hypothesis implying the joint hypothesis $(\tilde{\beta}, \tilde{\alpha}) = (0, 1)$ (Nelson and Plosser, 1982, p. 144).

There is some disagreement in the literature as to the order of hypothesis testing for Equations 2 to 4. Dickey *et al.* (1986) believe that testing should begin with Equation 3. Such statistics would have higher power if Equation 3 is the valid model, against statistics generated with Equation 2 or Equation 4. Perron (1988) disagrees and recommends starting with Equation 4. This allows a test of the unit root hypothesis against the obvious alternative that the series is trend-stationary. Under this alternative, Equation 3 will not be able to distinguish a unit root from a trend-stationary process. We choose to use Perron's strategy, and test the

[4]The ACF for $\Delta\Delta LGD$ has no significant autocorrelations at any lag.

unit root hypothesis directly against the trend–stationary alternative.[5]

In terms of Equation 4, the nested hypotheses are:

(1) H_0: $(\mu, \beta, \alpha) = (0, 0, 1)$. The Dickey and Fuller (1981) ϕ_2 test is computed with critical values to be found in their Table V.2, p. 1063. To account for the possibility that the errors in Equation 4 are not iid, $N(0, \sigma^2)$, as is assumed by Dickey and Fuller, we compute Phillips and Perron's (1986) $Z(\phi_2)$ statistic which is a modified version of the ϕ_2 statistic;

(2) H_0: $(\mu, \beta, \alpha) = (\mu, 0, 1)$. The ϕ_3 test statistic (Dickey and Fuller, 1981) is computed with critical values appearing in their Table V.3, p. 1063. The corresponding $Z(\phi_3)$ due to Phillips and Perron (1986) is also computed;

(3) H_0: $\alpha = 1$ in Equation 2. By imposing zero restrictions on μ, β and the higher order differenced terms, Equation 4 simplifies to $y_t = \alpha y_{t-1} + e_t$ which is the basis of the Dickey–Fuller test (DF). The test statistic is calculated as $(\alpha_1 - 1)$ divided by its standard error. Adding the higher order differenced terms provides the augmented Dickey–Fuller (ADF) test regression.

The test results are reported in Tables 2 and 3. Using the ϕ_2 test of the joint hypothesis of a unit root with no drift, we cannot reject the null for any of the variables in level form. The information obtained from the Phillips–Perron $Z(\phi_2)$ modification to the ϕ_2 test essentially confirms this conclusion, although at some lags the significance of the test is above the 5% level. In first difference form, the null can be rejected for all cases, except ΔLGD, although it can also be rejected for $\Delta\Delta LGD$. The $Z(\phi_2)$ tests for differenced variables (not reported but available on request) are supportive.

The null of a random walk with fixed drift (ϕ_3 test) cannot be rejected for any of the variables in level form. Again the $Z(\phi_3)$ tests are generally consistent with this result, although at lower lags the null for *LSIG* is only valid at the 2.5% level. For first differences, the null is rejected in every case bar ΔLGD. Again the hypothesis is rejected for the variable $\Delta\Delta LGD$.

The DF and ADF results, given our rejection of model 4, cannot reject the unit root hypothesis for all variables in level form and ΔLGD. The unit root hypothesis is rejected for the first-differences of *LUR*, *LCU* and *LSIG*, and the second-difference of *LGD*.

We tentatively conclude that *LUR*, *LCU* and *LSIG* are $I(1)$ and *LGD* is $I(2)$. As a result, the first difference of potential output, ΔLGD and the other variables in level form could constitute a cointegrating set.

Testing for cointegration

The CR can be estimated in a number of different inversions which produce different long-run parameters.[6] Equation 6 reports a cointegral relationship between the unemployment rate (as the dependent variable) and the level of capacity utilization, the first change of potential output, and

Table 2. *Testing for unit roots in Okun variables*

Variable	X(0)	T	DF	ADF	Φ_2	Φ_3
LUR	67(4)	87	+0.79	+0.20	2.10	2.76
ΔLUR	68(1)	86	−7.93	−4.35	7.56	11.34
LCU	67(4)	87	−1.11	−0.53	2.35	3.48
ΔLCU	68(1)	86	−14.49	−4.28	6.53	9.76
LGD	67(4)	87	+19.35	+1.47	4.06	4.87
ΔLGD	68(1)	86	−1.98	−1.87	1.96	2.39
$\Delta\Delta LGD$	68(2)	85	−8.94	−3.74	5.45	8.17
LSIG	68(1)	86	−0.76	−1.01	4.25	5.97
$\Delta LSIG$	68(2)	85	−14.86	−6.99	16.96	25.43

Notes: The critical values at the 5% level are 3.5 (DF), 3.2 (ADF) (Hall, 1986), and 4.88 (Φ_2) and 6.49 (Φ_3).

[5]The regression format employed is (Nelson and Plosser, 1982, p. 150)

$$y_t = \mu + \beta t + \alpha_1 y_{t-1} + \sum_{i=1}^{k} \alpha_{i+1} \Delta y_{t-i} + e_t$$

It is possible to employ an Akaike information criterion to test for the optimal number of higher order terms (k), but this approach did not prove fruitful. Guidance in this regard was principally derived from information contained in the ACFs (Table 1), and the regression t statistics in the higher order terms for different values of k. We use four lags.

[6]The use of *LUR* as the dependent variable in the cointegrating equation is optional. Using the other variables as the dependent variable would have generated different OLS results, and thus different estimated equilibrium relations (Hall and Henry, 1986). Following Henry and Hall (1988) we might interpret the variations in the results as defining the limits of the equilibrium space, with the exact long-run position being enclosed by the set so defined.

Table 3. *Phillips–Perron $Z(\phi_2)$, $Z(\phi_3)$ tests, 1967(1)–89(2): model 3*

	Truncation Lag					
LUR	2	4	6	8	10	12
$Z(\phi_2)$	1.91	1.75	1.98	2.28	2.39	2.20
$Z(\phi_3)$	2.86	2.62	2.97	3.41	3.59	3.30
LCU	2	4	6	8	10	12
$Z(\phi_2)$	2.58	2.52	2.45	2.57	2.77	3.01
$Z(\phi_3)$	3.86	3.78	3.67	3.84	4.12	4.45
LSIG	2	4	6	8	10	12
$Z(\phi_2)$	4.88	5.02	4.92	4.50	4.32	4.06
$Z(\phi_3)$	7.32	7.52	7.38	6.75	6.48	6.09
LGD	2	4	6	8	10	12
$Z(\phi_2)$	4.43	4.56	4.68	4.85	5.04	5.65
$Z(\phi_3)$	5.89	6.05	6.22	6.44	6.70	7.51
Critical values ($T=100$)			10%	5%	2.5%	1%
$Z(\phi_2)$			4.16	4.88	5.59	6.59
$Z(\phi_3)$			5.47	6.49	7.44	8.73

Source: Dickey and Fuller (1981).

the structural change variable. Although the ADF statistic is slightly low, the DF and CRDW statistics are significant.

For the period 1966(4)–1989(2), a total of 91 observations, the CR is:[7]

$$LUR = 4.16 - 4.54LCU - 53.43\Delta LGD + 0.58LSIG \quad (6)$$

$$R^2 = 0.74 \quad \text{CRDW} = 0.91 \quad \text{ADF} = 2.91 \quad \text{DF} = 5.11$$

The estimates in Equation 6 are biased but consistent and converge on their true parameter values more quickly than if the variables were stationary (Stock, 1987). Importantly, the extent of the small-sample bias is related to $(1 - R^2)$ of Equation 6, which suggests that in our case the bias is not large (Banerjee *et al.*, 1986). So despite the lack of dynamics in the CR and the resulting mis-specification, it is still possible to derive asymptotic estimates of the long-run multipliers, and hence an estimate of the steady-state rate of unemployment, conditional on the other variables.

IV. TESTING DOWN TO A DYNAMIC REPRESENTATION

The general dynamic model expresses the one-quarter change in the rate of unemployment (ΔLUR) as a function of its own lags, current and lagged values of the one-quarter change in capacity utilization (ΔLCU), the change in potential output (capacity) growth ($\Delta\Delta LGD$), and the change in the structural index ($\Delta LSIG$). Five lags of all variables are included. In addition, the one-quarter lagged value of the residuals from the CR, ECM(-1) is included to model the error correction component of the model. The largest available sample, 1968(2)–1989(2) is utilized and all variables are seasonally adjusted.

Initial OLS estimation of the model revealed no problems of residual autocorrelation, although two observations (1970(1) and 1973(3)) were associated with large residuals. To overcome problems of pretesting, increasingly stringent testing was used when valid simplifying restrictions were sought. A relatively simple search process resulted in the following parsimonious specification for the period 1968(2) to 1989(2):[8]

$$\begin{aligned}\Delta_1 LUR = {} & \underset{(2.53)}{0.019} - \underset{(5.74)}{0.136}ECM(-1) - \underset{(3.59)}{0.253}D701 \\ & - \underset{(3.60)}{0.248}D733 + \underset{(2.94)}{0.182}\Delta_1\Delta_1 LUR(-4) \\ & - \underset{(3.02)}{1.305}\Delta_1 LCU + \underset{(3.39)}{16.374}\Delta_2\Delta_1\Delta_1 LGD(-1) \end{aligned} \quad (7)$$

$$R^2 = 0.56 \qquad \text{RSS} = 0.3544 \qquad \sigma = 0.0674$$

[7]The variables in Equation 6 do not form a cointegrating set when the log of the real wage rate, an $I(1)$ variable, is included.

[8]The difference terms can be defined in general as:

$$\Delta_i\Delta_j LUR = (LUR - LUR(-j)) - (LUR(-i) - LUR(-i-j))$$

The diagnostics for Equation 7 are as follows:[9]

(A) $\chi^2(1)=0.29$ (B) $\chi^2(4)=0.49$ (C) $\chi^2(1)=3.05$
(D) $\chi^2(2)=0.64$ (E) $\chi^2(1)=0.06$

The restrictions on the general model which resulted in Equation 7 are all valid ($F[20, 58]=1.08$). The diagnostic test statistics reported show no evidence of functional form mis-specification, no first and higher-order serial correlation, and no problems of non-constant residual variance. ARCH tests were performed for lags one to eight inclusive, and all were satisfactory.

An examination of *CUSUM* and *CUSUMSQ* graphs for the period reveals no problems of within-sample instability. Most time series data of economic variables display considerable variability during the mid 1970s as a result of a series of large shocks (OPEC etc.) that the Western economies faced. As noted, the general model has two observations with large residuals (1970(1) and 1973(3)), which could be evidence of either one-off, abnormal behaviour which was subsequently absorbed by the system, or alternatively, complete structural breaks in the behavioural relations. Formal testing was performed to determine which alternative was more plausible. A Chow test was inappropriate to check for a structural break at 1970(1) because there were too few degrees of freedom in the first sub-sample. A dummy variable approach was used (*D70*=0 prior to 1970 and unity thereafter) where *D70* was included in intercept and interactive form. None of the variables (*D70* or the interactive terms) were significant.

The null that the coefficients are invariant between the split samples, 1968(2) to 1973(2) and 1973(3) to 1989(2) is conditional on the residuals being homoscedastic for the whole sample. Thus the correct sequence is to test the uniformity of residuals across both samples, before calculating the Chow statistic. The heteroscedasticity diagnostics (Test (D) above and the ARCH tests) left us satisfied that the error variances were uniform across the sample. The Chow test statistic for parameter stability, $\chi^2(6)=4.48$, supports the null that no significant change in the regression surface occurred at 1973(3). Similar results were obtained for various split samples in this neighbourhood. Post-sample stability was assessed using Chow's second stability test. The predictive failure statistics for four and eight quarters out of sample are both satisfactory ($\chi^2(4)=2.81$; $\chi^2(8)=4.20$).

While the dynamics in Equation 7 are somewhat complicated, the EC variable is highly significant, and its coefficient indicates that about 14% of the change in *LUR* per quarter is attributed to the disequilibrium between the actual unemployment rate and the steady-state rate of unemployment. Thus, the quarterly change in the unemployment rate is positive (negative), other things equal, whenever the actual level is below (above) the steady-state level.

To further test the EC representation, we relaxed the restrictions imposed on $ECM(-1)$ by virtue of the CR, Equation 6. This was achieved by excluding this term and including the CR variables in lagged level form within the final tested-down form of the dynamic model. The additional variables were therefore $LUR(-1)$, $LCU(-1)$ $\Delta LGD(-1)$, and $LSIG(-1)$. The results are shown in Equation 8:

$$\begin{aligned}\Delta LUR = & \underset{(3.75)}{0.518} - \underset{(3.27)}{0.231}D701 - \underset{(3.75)}{0.258}D733 \\ & + \underset{(2.78)}{0.179}\Delta_1\Delta_1 LUR(-4) - \underset{(3.75)}{1.751}\Delta_1 LCU \\ & + \underset{(3.54)}{16.834}\Delta_2\Delta_1\Delta_1 LGD(-1) - \underset{(6.02)}{0.143}LUR(-1) \\ & - \underset{(3.61)}{1.577}LCU(-1) + \underset{(2.96)}{0.076}LSIG(-1) \\ & - \underset{(2.99)}{7.310}\Delta_1 LGD(-1)\end{aligned} \tag{8}$$

$R^2=0.59$ $\sigma=0.0662$

(A) $\chi^2(1)=0.29$ (B) $\chi^2(4)=5.85$ (C) $\chi^2(1)=1.23$
(D) $\chi^2(2)=0.78$ (E) $\chi^2(1)=0.38$

The coefficient estimates of the dynamic variables are very similar to the two-stage model in Equation 7. The diagnostic performance of Equation 8 is also satisfactory.

The static long-run solution which is implied by Equation 8 is (normalizing on *LUR*):

$$LUR = 3.49 - 11.02LCU - 51.12\Delta LGD + 0.53LSIG \tag{9}$$

Equation 9 presents an equilibrium outcome similar to the CR, except for the higher coefficient on *LCU*. Further, all the lagged level terms are significant (at the 5% level) which is consistent again with the robustness of the EC representation in Equation 7. A joint test of zero coefficients on the level terms in Equation 9 provides further support for the validity of the EC hypothesis, with a test statistic ($F[4, 74]=9.87$) which is greater than the critical value at the 5% level. The advantage of Equation 7 is that the EC variable reduces the number of parameters in the model by three.

[9](A) $\chi^2(1)$=LM test of first order residual serial correlation.
(B) $\chi^2(4)$=LM test of fourth order residual serial correlation.
(C) $\chi^2(1)$=Ramsay RESET test using the square of the fitted values.
(D) $\chi^2(2)$=normality test based on skewed and kurtosis of residuals.
(E) $\chi^2(4)$=LM test for heteroscedasticity based on the regression of squared values on squared fitted values.
(F) $\chi^2(n)$=Chow's second test for predictive failure, where n is the number of observations omitted.

V. ANALYSIS

The macroequilibrium unemployment rate

It is possible to estimate the potential or steady-state level of unemployment, corresponding to different time periods from the CR. These estimates are shown in Table 4. The average unemployment rate for each time period is compared to the $U^*(1)$ which is calculated on the basis that the right hand variables in the CR ($LSIG$, ΔLGD, LCU) take their arithmetic means over the whole period 1966(4)–1989(2) and the subperiods 1966(4)–1973(2) and 1973(3)–1989(2). This yields steady-state values of the unemployment rate of approximately 4.65%, 2.15%, and 6.24%, respectively. $U^*(2)$ is based on the assumption that capacity utilization was 100% (that is, potential and actual output were always equal). The difference between $U^*(1)$ and $U^*(2)$ reflects the average cyclical variation in capacity usage in each time period.

The final subperiod (1984(1)–1989(2)) represents a period of sustained recovery and higher rates of capacity utilization, following a period of rising unemployment which culminated with a rate of 10.7% in February 1983. For this period, $U^*(1)$ is 6.26%, compared with a rate consistent with 100% capacity utilization ($U^*(2)$) of 4.64%.[10]

Other studies have found that in most countries the unemployment rate variable is stationary, $I(0)$, so that the rate of unemployment is mean reverting, which is consistent with the natural rate hypothesis. Our work shows that for Australia, the unemployment rate is $I(1)$, so that the unemployment rate exhibits hysteresis (Franz, 1990). In addition to the short and long-run cyclical variables, a structural change variable appears in the CR.[11]

Tables 5 and 6 decompose the estimates of the steady-state unemployment rate into three components. The short-run cyclical component assumes that ΔLGD and $LSIG$ are set at their average values corresponding to the period 1966(4) to 1973(2) for Table 5 and 1966(4) to 1989(2) for Table 6, whereas LCU takes its average value for each specific time period in both tables. Similarly, the structural variation component is based on the assumption that LCU and ΔLGD are set at their average values corresponding to the period 1966(4) to 1973(2) for Table 5 and 1966(4) to 1989(3) for Table 6, whereas $LSIG$ takes its average value for each specific time period in both tables. The capacity slowdown component is constructed in a similar manner.

The resulting steady-state unemployment rate estimates calculated from the CR using the above assumptions help us to understand which factors have influenced their movements over the different periods. With structural variation

Table 4. *Estimates of steady-state unemployment rates*

Time period	AVE UR	$U^*(1)$	$U^*(2)$
1966(4)–1989(2)	5.16	4.65	3.52
1966(4)–1973(2)	1.97	2.15	1.80
1973(3)–1989(2)	6.50	6.24	4.52
1984(1)–1989(2)	7.92	6.26	4.64

Notes: $U^*(1)$ refers to the estimate of U from the CR if all regressors take their average values over the relevant sample, whereas $U^*(2)$ assumes that potential and actual output are always equal.

Table 5. *Decomposing the estimates of steady-state unemployment rates*

Time period	Short-run cycle $U^*(3)$	Structural variation $U^*(4)$	Capacity slowdown $U^*(5)$
1966(4)–1989(2)	2.37	2.84	3.12
1966(4)–1973(2)	2.15	2.10	2.15
1973(3)–1989(2)	2.48	3.09	3.66
1984(1)–1989(2)	2.42	3.35	3.47

Notes: $U^*(3)$ assumes that ΔLGD and $LSIGSA$ are constant at their average values for the period 1966(4) to 1973(2), while LCU takes the actual average value for the relevant time period. $U^*(4)$ holds LCU and ΔLGD constant at average 1966(4) to 1973(2) values and $LSIGSA$ takes its own average value for the relevant time period. $U^*(5)$ holds LCU and $LSIGSA$ at average 1966(4) to 1973(2) values and ΔLGD takes its own average value for the relevant period.

Table 6. *Decomposing the estimates of steady-state unemployment rates*

Time period	Short-run cycle $U^*(6)$	Structural variation $U^*(7)$	Capacity slowdown $U^*(8)$
1966(4)–1973(2)	4.20	3.19	3.45
1973(3)–1989(2)	4.86	5.45	5.08
1984(1)–1989(2)	4.75	5.17	5.49

Notes: $U^*(6)$ assumes that ΔLGD and $LSIGSA$ are constant at their average values for the period 1966(4) to 1989(3), while LCU takes the actual average value for the relevant time period. $U^*(7)$ holds LCU and ΔLGD constant at average 1966(4) to 1989(3) values and $LSIGSA$ takes its own average value for the relevant time period. $U^*(8)$ holds LCU and $LSIGSA$ at average 1966(4) to 1989(3) values and ΔLGD takes its own average value for the relevant period.

[10]The estimated steady-state values corresponding to the CR, derived from the lagged levels regression 9, for each time period are close to those from the estimated CR, which again provides support for the robustness of the EC mechanism (for example, for the period 1966(4) to 1989(2), $U^*(1)$ was 4.43 and $U^*(2)$ was 2.26).

[11]Hysteresis has also been detected in an Australian wage inflation equation by Coe and Gagliardi (1985), Mitchell (1987), Coe (1988) and Watts and Mitchell (1990).

and capacity growth kept constant at their average values, the estimates of the equilibrium unemployment rate, $U^*(3)$, $U^*(6)$ show that short-run variations in capacity utilization are only responsible for modest fluctuations in the steady-state unemployment rate.

On the other hand, there are relatively large fluctuations in the steady-state unemployment rate caused by variations in both the structural variable and the long-run cyclical variable, the growth of capacity.

The index of structural change, *LSIG*, is a proxy for overall labour market imbalance. In Australia and most Western countries over the last 20 years, there has been a significant change in the composition of employment by industry and occupation. The growth of female-dominated subordinate, white collar occupations (e.g. clerical and sales) has been accompanied by an absolute decline in employment in skilled, male-dominated, blue collar occupations (Watts and Rich, 1991). The female and part-time shares have increased at the expense of male and full-time shares.[12] The labour force participation behaviour of the sexes differs markedly. Women exhibit a strong discouraged-worker effect, so that a rise in the employment:population ratio of working age women has had little impact on their unemployment:population ratio (Gregory, 1991).[13] By contrast, the displacement of male employees has been accompanied by a rise in their unemployment ratio, although, over time, the male participation rate has declined (Watts and Rich, 1991). Hence the structural change variable, which is measured by the variance of employment growth rates by industry, also captures the reinforcing impact of changes in supply behaviour of males and females.

The disturbing message from the arithmetic exercise is that, while hysteresis effects can reduce the impact of structural imbalance to a small degree, the rate of capacity growth places a floor in the full employment/unemployment rate (*FNUR*). A slowdown of capacity growth leads to capacity-constrained unemployment (Malinvaud, 1980), so that factors which retard investment in productive activity will constrain the minimum attainable *FNUR*. On the other hand a large variance of employment growth promotes a structural mismatch between the unemployed and vacancies, which takes time to resolve.

Standard Okun's Law studies which regress (in some form) the rate of unemployment on the output gap obscure the long run impact on unemployment of variations in the structural index of employment growth variance and capacity growth.

The Okun coefficient

Given the logarithmic specification, the Okun coefficient is non-constant but can be established for any combination of the rate of capacity utilization, CU, and the steady-state rate of unemployment.[14] From Equation 1, the Okun coefficient, β, can be written as

$$\beta = 1/q = \mathrm{d}YGAP/\mathrm{d}U = 100/(\varepsilon.\mathrm{CU.U}) \qquad (10)$$

where ε denotes the elasticity of the unemployment rate with respect to the rate of capacity utilization estimated from the CR.

We calculate Okun coefficients for the four periods using the constant unemployment/capacity utilization elasticity derived from the CR, the corresponding mean values of the CU and the steady-state unemployment rate, $U^*(1)$. The Okun coefficients are 5.04 for 1966(4)–1989(2), 10.66 for 1966(4)–1973(2), 3.79 for 1973(3)–1989(2), and 3.76 for 1984(1)–1989(2). These estimates are somewhat higher than those quoted by Nguyen and Siriwardana (1988).

With capacity utilization constant, our model shows that the Okun coefficient is inversely related to the steady-state unemployment rate. This means that in the absence of significant labour hoarding, recovery from a deep recession is accompanied by a significant cut in the unemployment rate. In periods of high pressure, a given fall in the output gap is accompanied by a smaller cut in the unemployment rate.

This is consistent with the hysteresis story. At low rates of unemployment, a person will either be frictionally unemployed or be in the most obdurate category of structural unemployment. A simple demand stimulus, in the absence of enhanced information aimed at reducing frictions and training/re-location/anti-discrimination programs geared to the long-term unemployed, is unlikely to achieve a comparable cut in unemployment as would occur at high unemployment. On the other hand, when there is low pressure in the economy, many of the unemployed appear to be structurally unemployed because they have lost or have been unable to obtain relevant job skills. As pressure rises, on-the-job training opportunities arise as vacancies are posted, and the structural imbalance diminishes (Mitchell, 1987). Employers would tend to seek other forms of adjustment, such as increasing hours of work for existing employed workers.

VI. CONCLUSION

Previous econometric approaches to the establishment of an Okun's law relationship have been based on inappropriate econometric procedures, in that relationships between non-stationary variables have been estimated.

Further, over the period in question, Okun's law is a misnomer, in the sense that other variables, namely a proxy for structural change and a long-run cyclical variable, influence the long-run determination of the unemployment rate.

[12]There is limited substitutability between these categories of employment (Gregory, 1991; Watts and Rich, 1991).
[13]A rise in the female part-time employment ratio appears to lead to a rise in their unemployment ratio (Gregory, 1991).
[14]By contrast, in the linear specification of Nguyen and Siriwardana (1988), for example, the Okun coefficient is constant.

In Australia there has been a long-run decline in the rate of capacity growth, which has been accompanied by an increase in the structural imbalance variable. Thus an immutable Okun's law relationship does not exist and a constant Okun coefficient over this period is denied.

The presence of a unit root in the unemployment rate reveals that it exhibits hysteresis, so that non-transitory changes in the equilibrium unemployment rate are caused by cyclical as well as other variables.

Finally, the re-specified equilibrium unemployment relationship casts some light on the demise of the Phillips curve in Australia. In an econometric analysis of Australian wage inflation, Watts and Mitchell (1990) demonstrate that a firm-based measure of pressure, namely the change in capacity utilization, is the appropriate excess demand proxy in the determination of wage inflation, as opposed to an aggregate labour market variable, such as the unemployment rate.

The unemployment rate would continue to explain wage outcomes as long as unemployment and capacity utilization exhibited a stable relationship. This occurred in Australia in the form of the Okun's law relationship until the early 1970s when the relationship broke down. Consequently, the Phillips curve relationship broke down at about the same time.

APPENDIX

All the time series variables are from the NIFJUN89 Data Series available on disk from the Australian Bureau of Statistics, except for Gross Non-Farm Product (*GNFP*), valued at 1984/5 prices, which is obtained from the National Accounts.

Potential capacity and capacity utilization variables

NIFJUN89 produce potential output and capacity utilization series, namely *GNMP* and *GUT*, based on the peak-to-peak approach. The output series employed, however, is a seasonally adjusted non-farm (constant price) series, *GNM*. The chosen peaks therefore do not coincide with the actual peaks but instead are statistical artefacts based on the chosen form of seasonal adjustment. It would seem to be more appropriate to construct the peak-to-peak potential series by using a seasonally unadjusted output series.

A further problem associated with the *GNMP* series is that peaks are chosen in such a way as to impose a declining trend on the series. This procedure is not only arbitrary but may entail substantial revision in the event of lengthening or revising the data due to the deletion of one or more intervening peaks (Watts and Mitchell, 1988).

A more suitable peak-to-peak approach is to adopt a more stringent criterion for the identification of local peaks which is only sensitive to data revisions, and is not sensitive to the lengthening of the sample. The criterion employed in the construction of the potential output, *GD* series is that a local peak must exhibit dominance, in the sense of exceeding in magnitude the adjacent four observations. For the final peak, dominance of the previous observations is the requirement. Linear, rather than exponential growth of potential output links consecutive peaks. If required, the prevailing growth rate is used to extrapolate beyond the final peak. Under this requirement, successive growth rates of capacity (potential output) are not required to decline.

Using the peak dominance criterion for the seasonally unadjusted series *GNFP*, local peaks are identified at 1973(4), 1976(4), 1981(4) and 1988(4). The first observation 1966(3) is used as the first peak. Kalisch (1982), cited in Nguyen and Siriwardana (1988, p. 19) runs a linear trend through quarterly observations of *GNM* over the period (1966–1971(4)), claiming that this epoch was a golden era of full employment, so that all unemployment was frictional.

This procedure for calculating capacity output is highly unsatisfactory because it implies a uniform absolute increase in capacity, quarter by quarter, and hence a slowing down of the rate of capacity growth as capacity expands. Also, it has been shown that there was a discrete shift in the trend rate of *GNM* growth in the early 1970s.

Derivation of the structural change index

Quarterly employment by industry data is only available in Australia from August 1978. Prior to that only annual (August) observations were published. BLMR (1987) derive a quarterly series by linearly interpolating between the August observations for years prior to 1978. The resulting series is smooth from 1966(3) to 1978(3), and then exhibits considerable fluctuation after 1978(3). To overcome this dichotomy, a general autoregression was fitted to each industry employment series for the annual observations from August 1966 to August 1989. Quarterly data for each industry was then derived from simulations of the autoregressions. Structural change is proxied by the Stoikov index, namely the standard deviation of annual growth rates of industry employment, SIG which is written as:

$$[\Sigma_i (N_{it}/N_t)(\mathrm{d}LN_{it} - \mathrm{d}LN_t)^2]^{1/2}$$

where N_t is employment in the *i*th industry and N is total employment.

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