AUSTRALIAN WAGE INFLATION: REAL WAGE RESISTANCE, HYSTERESIS AND INCOMES POLICY: 1968(3)–1988(3)*

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I   INTRODUCTION

Australia is unique among English-speaking countries in that, since 1975, it has persisted with different forms of incomes policy under the auspices of the Arbitration Commission, despite the general push in these countries towards more flexible wage-fixing arrangements. Yet, economists have claimed that incomes policies have been ineffective in the long run in Canada, Britain and the U.S.A. (see Christofides and Wilton, 1985, p. 51, for relevant references).[1] In Australia, the period 1968(3)–1988(3) is a fertile era for empirical enquiry into the effects of incomes policy on wage outcomes because there were four broadly distinct regimes of wage fixation, namely decentralized bargaining (1968(3)–1975(1), 1981(2)–1982(4)), full/plateau/partial indexation (1975(2)–1981(2)), a wages pause (1983(1)–1983(2)) and the Accord Phase 1 (1983(3)–1987(1)) and the Accord Phase 2 (1987(2)–1988(3)). In short, a *de facto* incomes policy has operated in Australia since 1975(2),[2] except for a brief return to collective bargaining in the early 1980s. Table 1 outlines the main phases of wage-setting principles since 1975.

[1] The moderation of wage inflation during the lifetime of the controls tends to be offset by catch-up once the controls are removed.

[2] The 1975(2)–1981(2) phase can be divided into four sub-periods corresponding to variations in the guidelines used. The unifying aspect of the incomes policy in this period though was the fact that the Arbitration Commission was concerned with the degree of money wage indexation as long as substantial compliance to the wage fixing guidelines was observed by all parties. The wages pause outlawed any increase of money wages. Under the Accord, both money wage adjustments (via indexation) and real wage movements (by a formal productivity award) were to be decided upon through formal National Wage Case (NWC) hearings under the direct jurisdiction of the Arbitration Commission.

TABLE 1
INCOMES POLICY PHASES

Wage-Setting Principles		Dates
Mark 1 Guidelines	(*IP*1)	1975(2)–1976(2)
Mark 2 Guidelines	(*IP*2)	1976(3)–1978(2)
Mark 3 Guidelines	(*IP*3)	1978(3)–1979(3)
Mark 4 Guidelines	(*IP*4)	1979(4)–1981(2)
Wages Pause	(*IP*5)	1983(1)–1983(2)
Accord Phase 1	(*IP*6)	1983(3)–1987(1)
Accord Phase 2	(*IP*7)	1987(2)–1988(3)

Recent Australian studies of wage inflation reveal that the (Lipsey–Parkin) Phillips curve specification is subject to instability (e.g., Gregory, 1986, and Dawkins and Wooden, 1985). An international study by Coe and Gagliardi (1985) and Mitchell's Australian work (1987) provide evidence of hysteresis of the unemployment rate. Under hysteresis the steady-state unemployment rate is cyclically sensitive. Consequently, the unemployment rate does not impinge directly on wage inflation, as in the constant natural rate scenario. Rather the difference between the current rate of unemployment and the steady-state rate of unemployment assumes importance where the latter is defined as a function of lagged values of the unemployment rate. Mitchell (1987) finds evidence of a stable aggregate wage inflation equation for Australia which is based on a specification incorporating real wage resistance, an aggregate hysteresis term and the effects of incomes policy in the form of award adjustment.

The impact on aggregate wage inflation of the most recent phase of incomes policy, namely the Accord, does not appear to have been subjected to detailed study and, in addition, a more thorough investigation of the resistance and hysteresis effects is warranted. Accordingly, in this article we attempt to establish whether, over the period 1968(3)–1988(3), the different phases of wage guidelines can be statistically identified.

In addition, we nest competing hypotheses about the impact of real wage variables and different forms of labour market pressure variables within a general specification of the wage inflation process. We use the *general to specific* method (Hendry, 1980), beginning with an over-parameterized model with liberal lags on all variables (excluding the dummy variables). The testing-down procedure continually simplifies until a satisfactory parsimonious representation of the data generating process (DGP) is achieved.

Apart from satisfying statistical criteria (Hendry, 1980), the model must be theoretically consistent, which usually refers to the compatibility between the short- and long-run properties of the estimated model. We generalize this to embrace the requirement that the preferred equation must yield a meaningful theoretical interpretation. The cointegration literature is relevant

here. While Box–Jenkins models overcome the spurious regression problem outlined by Granger and Newbold (1977), the specification in differenced form cannot capture possible long-run relations between levels. Error Correction Models (ECMs) which included both level and differenced variables in the general form were designed to overcome the problem of spurious regression yet model the long-run properties. Cointegration solves the problem of estimating ECMs if the level terms are not stationary. If the residuals are non-stationary, all conventional significance tests are invalid.

A two-step approach (Engle and Granger, 1987) to ECM estimation allegedly overcomes this problem with conventional ECM modelling. The long-run surface is first estimated and, if cointegration is established, the lagged residuals of this first stage can be used in the dynamic estimation as representative of the disequilibrium in the system. This technique ensures that the residuals in the second stage are stationary. Unfortunately, the power of the cointegration and related unit root tests is low and the two-step procedure is still *ad hoc*. Although the theoretical properties of cointegrated sets are important to dynamic modelling, in practical terms the results we get from conventional ECM modelling are superior to the two-step approach at this stage.

In Section II, the relevant literature is summarized. The model is outlined in the following section. Section IV presents the results of our dynamic model (conventionally estimated), while Section V introduces the concept of cointegration and the use of two-step error correction estimation due to Engle and Granger (1987). Conclusions are drawn in Section VI.

II   PREVIOUS STUDIES

A number of studies have explored the impact of incomes policies on earnings inflation in Australia (see Newton and Kalisch, 1983, for a recent survey of Australian studies of wage inflation in general). Phipps (1981); Gregory and Smith (1985); Gregory (1986); and Zaidi (1986) have all recently examined the impact of incomes policies on wage outcomes in Australia. Their results broadly point to some degree of earnings moderation during the operation of the wage-setting guidelines between 1975 and the early 1980s, although their findings are not uniform. In general, they demonstrate a failure to use any systematic econometric methodology with respect to model selection, being content to draw conclusions based on models with poor diagnostics without evidence of any extensive econometric testing.

The policy-off simulation method is used by Phipps (1981); Gregory (1986); and Zaidi (1986). Phipps (1981) models average weekly earnings inflation over the non-indexation period 1961(3)–1975(1). His reported Durbin–Watson statistic is very low, which must cast doubt on his estimates. The simulations of the equation over the indexation phase 1975(2)–1980(2), assuming strict exogeneity of all right-hand side variables, indicate that the

estimated wage equation, representing the free market, yielded a higher aggregate outcome between 1975(2) and 1977(2) but a lower rate of wage increase during the period of high unemployment—1977(3)–1980(2).

Gregory (1986), extends Phipps's sample to include the 1981(3)–1982(4) period which followed the post-1981 interregnum of collective bargaining. His average weekly earnings equation based on a standard Phillips curve (the unemployment rate being the excess demand proxy) is out-performed by a specification which uses overtime as the labour market pressure variable. The rationale is that a firm-level utilization variable is more relevant to wage inflation than an aggregate labour market measure due to employer-employee attachments and the notion that implicit contracts deliver real wage increases, independently of the state of the external labour market (s. 64). His simulations of the policy-off estimates reveal that indexation sustained a higher rate of wage increase over the first two years than otherwise, whereas, subsequently, wage outcomes were lower than predicted by a policy-off equation.

While each of these studies has some econometric difficulties which cast doubt on the robustness of their estimates, the main problem relates to the method employed to examine the effects of incomes policies on wage inflation. The validity of the policy-on, policy-off approach depends on the assumption that the behaviour of the relationship in the policy-off period is invariant to the incomes policy. Usually, no formal procedure is employed to test for the exogeneity of the regressors. Further, estimation of a policy-off sample with observations from the pre- and post-incomes policy period is also problematic (see Gregory, 1986, and Henry and Ormerod, 1978, p. 33) because incomes policies are alleged to generate pent-up real wage aspirations which are hastily realized in catch-ups once the policy is relaxed. In this case, the total policy-off sample will not reflect consistent behaviour. No tests are usually employed to detect this effect.

Zaidi (1986), presumably influenced by the work of Hagens and Russell (1985), examines the "intercept-shift hypothesis" and the "rotational approach", in addition to estimating and simulating a policy-off equation. He thus estimates, separately, three specifications: (a) a policy-off equation; (b) an equation with level dummies covering the various policy-on periods; and (c) an equation with interactive dummy variables to test the rotation hypothesis. Zaidi's intercept and rotation regressions do not prove successful. He concludes, after simulating the policy-off equation over the policy-on era, that the labour market variables consistently emerge as the restraining influence on wage inflation. The estimated Phillips curve equation has very limited explanatory power and poor diagnostic performance, so any conclusions drawn from it are unjustified.

A more correct and general approach is to estimate a full-sample equation (ignoring the arbitrary policy-off, policy-on dichotomy for reasons already mentioned) which nests the shift and rotation hypotheses and tests

down to a more specific form. The *IP* impacts are subsequently determined by the coefficient sign and magnitude on the relevant and significant variables.

The approach adopted in this article is an advance on the previous Australian work in three ways: (a) in the light of the previous discussion, a more valid econometric approach is used, in that we test down from a general specification of behaviour over the whole period; (b) we test whether the type of incomes policy is important by including variables for all the separate guidelines—including the indexation era of the Accord (Phase 1) and two quarters of the two-tier system (Phase 2); and (c) we examine the period (1981–1983) which is nested between periods of explicit incomes policy to test whether any catch-up effects can be identified.

III THE MODEL

The hypotheses which our model nests in the general specification include inertia of money wage growth; real wage resistance; a cyclically sensitive steady-state unemployment rate (the hysteresis effect), which generates a long-run inflation–unemployment trade-off (with homogeneity as a special case), *versus* the alternative excess demand theory (the Phillips curve) and the impact of the different incomes policy guidelines. The institutional features of the model are explained in Mitchell (1987) and Sargan (1964, 1980). All the equations in this article are in log-linear form and were estimated using ordinary least squares.[3]

The general form of our wage equation is as follows:

$$\Delta_4 AWE = \alpha_0 + \alpha_1(L)\Delta_4 AWE(-1) + \alpha_2(L)\Delta_4 CPI + \alpha_3(L)RW(-4)$$
$$+ \alpha_4(L)Z + \sum_{i=1}^{7} \beta_i IP_i \qquad (1)$$

where $\alpha_1(L)$, $\alpha_4(L)$ are fifth-order polynomials and $\alpha_2(L)$, $\alpha_3(L)$ are fourth-order polynomials. The fourth lag on the dependent variable was deleted to avoid problems with the real wage term. $\Delta_4 AWE$ denotes the four-quarter change of the log of average weekly earnings, $\Delta_4 CPI$ is the corresponding change of the log of the consumer price index, Z is the log of the variable measuring excess demand or bargaining power in the labour market (hereafter termed pressure variable), RW is the level of real average weekly earnings, and IP is a matrix of dummy variables with each column designed to account for a separate phase of centralized wage setting guidelines (0 = policy-off;

[3]A Wu-Hausman test for the independence of the capacity utilization term in the parsimonious specification, (1.2), namely $\Delta_1 GUT$ and the error term in the model (using $\Delta_4 GUT(-1)$, $\Delta_1 GUT(-4)$ and the other regressors from the model as instruments) yielded an F-statistic $[1,66] = 3\cdot52$. Thus we conclude that $\Delta_1 GUT$ is statistically exogenous and the OLS estimates will be unbiased.

1 = policy-on), identified in Table 1. The *IP* variables also enter the model interactively to test for slope sensitivity. A dummy variable *TD*1 (unity in 1974(3) and zero otherwise) is also included which is designed to capture the abnormal wage developments in that quarter identified by Gregory (1986). All the variables in the estimations are explained in detail in the Data Appendix.

The Dependent Variable—Average Weekly Earnings

To focus on unit labour costs and hence the price level, it would be natural to use the growth in earnings per hour as the dependent variable. This would also avoid the problem noted by Gregory (1986, s. 73) of spurious correlation between average weekly earnings and labour utilization rates within the firm. Typical studies of wage inflation, however, focus on the growth of average weekly earnings (see Newton and Kalisch, 1983; Gregory and Smith, 1985).[4]

This may follow from the possible difficulties encountered when earnings per hour is used. The dependent variable becomes a ratio of two variables, each of which may be positively correlated with excess demand pressures. Consequently, the sign of the pressure variable in an hourly earnings equation is ambiguous.

While the issue of the homogeneity of earnings with respect to hours worked is interesting in its own right (because it allows insights into the relative price and quantity adjustments that firms might employ as the business cycle evolves), the possible direct and indirect influences of variations in activity on inflation need to be more explicitly estimated. Accordingly, we begin with average weekly earnings as the dependent variable and then use average weekly hours in an added variable test to study the relative strengths of price and quantity adjustments.

Four-quarter log changes are preferred *a priori* on the grounds that this more adequately captures the successive wage and price adjustment patterns of the Australian wage-setting system. The claim that four-quarter changes introduce autocorrelation (see Kirby, 1981; Zaidi, 1986) is an econometric issue and should not necessarily guide, *a priori*, the appropriate specification. At the outset, the model should attempt to capture the flavour of the data generating process rather than accepting statistical expediency.

Hypotheses Nested in the General Specification

a) *Inertia of Money Wage Growth.*   Five lags of the dependent variable are employed in the general specification to capture the notion of inertia of the

[4]A more appropriate dependent variable would be average weekly ordinary time earnings which is free from the influence of hours variation, classification changes and sex composition changes. Unfortunately this series only begins in the mid-1970s.

money wage bargaining process. Inertia in the growth of money wages is the dynamic outcome of comparison effects or, in Australian terminology, *comparative wage justice*.

b) *Real Wage Resistance.*   Wage outcomes reflect the interaction of a host of contingent actions by individuals given the limited information about the uncertain future. These plans and their behavioural manifestations are, in turn, channelled into outcomes through the interplay of a range of institutional forces, like the Arbitration Commission and associated Tribunals, Trade Unions, Employer Groups and the Government. The plans involve the formation of expectations of future events and adaptive adjustments when these expectations are found wanting.

The econometric model should test for the equilibrium relations between the variables, in addition to capturing the tension of the dynamic adjustment induced by disequilibrium within the system. This implies that an appropriate specification should incorporate the possibility of error correction behaviour (see Hendry, 1980, pp. 232–233, and Wadhwani, 1985, p. 198). In Section V, the alternative cointegration approach to this problem is briefly explored.

Bargaining over nominal wages may be influenced by a concern to achieve some target real disposable earnings.[5] Following Sargan (1980); Henry and Ormerod (1978); and Wadhwani (1985), the form of real wage resistance (or error correction) can be examined within the general specification by the imposition of the appropriate restrictions on the lag operator polynomial $\alpha_3(L)$ attached to the real wage terms.

The most general hypothesis is that lagged real earnings affect the current wage outcome. We use the fourth lag of the real earnings variable as our starting observation because our model is based on annual changes (see Wadhwani, 1985, p. 198). This approach avoids the potential problem of collinearity of the regressors given the presence of lagged dependent variables and lagged inflation terms. Such a specification enables the empirical differentiation of nominal inertia and real wage resistance effects.

The simplest form of real wage resistance embraces a constant target rate of real wage growth (see Henry and Ormerod, 1978) but then the trend (constant) term becomes a catch-all term and it does not allow the revision of the target rate of real wage growth (see Wadhwani, 1985). The inclusion of the trend which is a non-stationary variable poses the problems identified in the cointegration literature. A more explicit form of the resistance hypothesis can be represented by the deviation of the fourth lag of the log of the real earnings

[5]Neoclassical theory (Parkin *et al.*, 1976) models excess demand for labour as a function of the real wage. Thus the movement in nominal wages depends on inflationary expectations and the extent to which the real wage deviates from equilibrium. The latter is considered an additional excess demand measure which helps to explain the poor performance of simple excess demand proxies.

variable from its endogenous target level which is defined as its lagged four-quarter moving average.[6]

c) *Aggregate Labour Market Pressure.*   Mitchell (1987) and Coe and Gagliardi (1985) both substantiate an hysteresis effect operating in Australia in the post-war period. Following Coe and Gagliardi (1985), the specification of the pressure variable MUG (macroeconomic unemployment gap) as the deviation of the actual unemployment rate log (U) from the steady-state rate (NRU) is a restricted form of the general lag operator polynomial $\alpha_5(L)$, if NRU is represented by a four-quarter lagged moving average of log (U).

If the restrictions on $\alpha_5(L)$ implied by the construction of the hysteresis variable are confirmed, then the vertical long-run Phillips curve is invalidated. Examining the validity of the restrictions provides a test of the hysteresis effect against an alternative cumulative impact effect. The ratio of vacancies to unemployment $(VONU)$ and the vacancies to labour force ratio (JVR) were also tried, one at a time, in the general specification.

d) *Capacity Utilization Pressure in Insider Labour Markets.*   According to the hysteresis hypothesis, rising aggregate unemployment "drives a wedge between the potential and actual labour supply and to some degree insulates the wage demands of the employed from the business cycle" (Mitchell, 1987, p. 104). Hysteresis could operate through capacity utilization and testing involves determining whether the utilization variable enters the simplified specification in level or differenced form.

The rate of labour utilization within individual firms and firms' expected profitability and access to credit are deemed to be the key to the capacity of employers to maintain money wage growth, rather than the state of the aggregate labour market (see Gregory, 1982, 1986). Further, organized labour is sensitive to utilization rates in that they provide the best guide to employment security (see Mitchell, 1987).

Thus a utilization variable may more adequately capture the bargaining by "insulated" workers. In this work, we utilize a range of within-firm aggregate utilization measures. First, the current and lagged values of the overtime rate (ROT) and average weekly hours (AWH) were tried. A second weekly hours series based on the deviation from its quadratic trend $(AWHD)$ was also constructed and tried. The national accounts-based utilization variable GUT (deviation from GDP from a linked-peaked GDP potential series) proved most satisfactory.

[6]In his work on U.K. wage inflation, Wadhwani (1985) successfully identifies resistance based on a 12-quarter moving average (target) real wage but the coefficient restrictions implied by the target real wage are not tested. Likewise, in our work, *a priori* restrictions on the RW term consistent with $RW(-4)$ and the four-quarter target variable test down successfully to the real wage resistance specification described in the text.

IV RESULTS

The empirical work covers the period 1968(3)–1988(3). Variables depicted as four-quarter changes in the estimated equation are seasonally unadjusted whereas, for consistency, the real wage variables and the excess demand proxies which appear in level form are seasonally adjusted. The most general specification used is represented in equation (1) with the pressure variables, described above, considered one at a time. All the incomes policies variables in level and interactive form are included, although the latter were not significant and were eliminated.

None of the aggregate pressure variables proved satisfactory. The diagnostic performance of all general equations using the aggregate pressure variables was poor, indicating serious mis-specification, with strong evidence of serial correlation. A simplified model testing the presence of the aggregate hysteresis hypothesis with respect to the unemployment rate accepted the MUG restrictions at the 5 per cent level, although the hysteresis effect was not statistically significant.

Both $VONU$ and JVR remained statistically significant in a simplified equation which incorporated money wage inertia, real wage resistance and hysteresis. However, in all the above-mentioned cases, the simplified forms are illegitimate due to the presence of serially correlated residuals and functional form mis-specification, which negates the validity of the F-tests used to simplify the general specification.

Having rejected the role of aggregate (labour market) pressure effects, attention was focused on the specifications using utilization pressure variables. Equations with the hours variables AWH and $AWHD$ and overtime ROT were found to be unsatisfactory. The GUT measure generated superior results.

The general equation is reported below as equation (1.1).

$$\Delta_4 AWE = -0.06 - 0.02 IP1 - 0.03 IP2 - 0.02 IP3 - 0.02 IP4$$
$$\quad (0.38)\quad (1.56)\qquad (3.52)\qquad (2.67)\qquad (2.03)$$

$$-0.04 IP5 - 0.03 IP6 - 0.03 IP7 + 0.90 \Delta_4 AWE(-1)$$
$$\quad (3.68)\qquad (3.06)\qquad (2.70)\qquad (7.83)$$

$$-0.04 \Delta_4 AWE(-2) - 0.13 \Delta_4 AWE(-3)$$
$$\quad (0.30)\qquad\qquad (0.90)$$

$$+0.11 \Delta_4 AWE(-5) + 0.52 \Delta_4 CPI - 0.59 \Delta_4 CPI(-1)$$
$$\quad (1.10)\qquad\qquad (3.10)\qquad\quad (2.82)$$

$$+0.45 \Delta_4 CPI(-2) - 0.09 \Delta_4 CPI(-3) - 0.07 \Delta_4 CPI(-4)$$
$$\quad (2.04)\qquad\qquad (0.42)\qquad\qquad (0.36)$$

$$+0.59 GUT - 0.65 GUT(-1) + 0.47 GUT(-2)$$
$$\quad (3.33)\qquad (3.43)\qquad\quad (2.40)$$

$$-0.31 GUT(-3) + 0.15 GUT(-4) - 0.24 GUT(-5)$$
$$\quad (1.48)\qquad\qquad (0.74)\qquad\qquad (1.45)$$

$$-0.55 RW(-4) + 0.44 RW(-5) + 0.08 RW(-6)$$
$$\quad (3.99)\qquad\quad (2.85)\qquad\quad (0.53)$$

$$-0.02 RW(-7) + 0.06 RW(-8) + 0.07 TD1$$
$$\quad (0.09)\qquad\quad (0.42)\qquad\quad (4.79)\qquad\qquad (1.1)$$

$$\bar{R}^2 = 0.94 \qquad \sigma = 0.01 \qquad RSS = 0.007 \qquad \text{Sample: } 1968(3)–1988(3)$$

(absolute t-values in parentheses).

An examination of equation (1.1) suggests the following simplification:

$$\Delta_4 AWE = \alpha_0 + \alpha_{10}\Delta_4 AWE(-1) + \alpha_{22}\Delta_4 CPI(-2) + \gamma_2 \Delta_1 \Delta_4 CPI$$
$$+ \gamma_3 \Delta_1 RW(-4) + \gamma_4 \Delta_1 GUT + \sum_{i=1}^{7} \beta_i IP_i \qquad (2)$$

with the following restrictions being imposed:

$$\alpha_{20} = -\alpha_{21} = \gamma_2$$
$$\alpha_{30} = -\alpha_{31} = \gamma_3$$
$$\alpha_{40} = -\alpha_{41} = \gamma_4$$

Equation (1.2) reports the results for the restricted version of (1.1).[7]

$$\Delta_4 AWE = +0.01 - 0.02 IP1 - 0.02 IP2 - 0.02 IP3 - 0.01 IP4$$
$$\quad (2.84)\quad (2.98)\qquad (3.67)\qquad (2.70)\qquad (2.14)$$

$$-0.05 IP5 - 0.02 IP6 - 0.02 IP7 + 0.84 \Delta_4 AWE(-1)$$
$$\quad (5.11)\qquad (4.42)\qquad (3.56)\qquad (15.82)$$

$$+0.63 \Delta_1 \Delta_4 CPI + 0.21 \Delta_4 CPI(-2) + 0.43 \Delta_1 GUT$$
$$\quad (4.79)\qquad\qquad (2.72)\qquad\qquad (3.04)$$

$$-0.66 \Delta_1 RW(-4) + 0.07 TD1$$
$$\quad (6.96)\qquad\qquad (5.19)\qquad\qquad\qquad (1.2)$$

$$\bar{R}^2 = 0.94 \qquad \sigma = 0.01 \qquad RSS = 0.010 \qquad \text{Sample: } 1968(3)–1988(3)$$
$$\chi_1^2(1) = 2.78 \qquad \chi_2^2(4) = 9.86 \qquad \chi_3^2(1) = 0.06 \qquad \chi_4^2(1) = 0.77$$

(absolute t-values in parentheses).

[7] $\chi_1^2(1)$ is a Lagrange multiplier test of first-order serial correlation; $\chi_2^2(4)$ is a Lagrange multiplier test for fourth-order serial correlation; $\chi_3^2(1)$ is the Ramsey RESET test statistic based on the square of the fitted values; and $\chi_4^2(1)$ is a Lagrange multiplier test statistic derived from a regression of squared residuals on squared fitted values. The Data-FIT econometric modelling program was used throughout.

The restrictions are strongly accepted by the data (F[15, 52] = 1·41), and thus, (1.2) is a more than reasonable simplification of the general form (1.1). The expected signs on the regressors are validated and the magnitudes of the coefficients appear plausible.

Diagnostically, the restricted equation performs very well, exhibiting no problems of functional form mis-specification, or heteroscedasticity. The LM test for fourth-order serial correlation is marginal (at the 5 per cent level), although the modified (F[4, 63] = 2·18, clearly rejects the null (at the 5 per cent level). No residual exceeds in absolute value twice the standard error, and Dickey–Fuller and Augmented Dickey–Fuller tests reject the unit root hypothesis, suggesting that the residuals are stationary.

Graphical analysis based on recursive residuals was conducted (CUSUM and CUSUMSQ) with no evidence of instability found. The LM test for uniform error variance ($\chi_3^2(1) = 0.77$), and ARCH tests for various lags were satisfactory. As such, it is valid to use Chow tests to explore the null of within-sample parameter constancy. Two split sample tests were conducted (coinciding with relatively large residuals), the first split at 1974(2) which yielded a F[6, 69] = 1·28; while the second at 1983(2) generated an F[6, 69] = 1·39. It is reasonable to conclude that (1.2) exhibits no evidence of within-sample instability.

The sample was shortened by 5 observations (to avoid singularity problems due to $IP7$). Chow's F-test statistic for predictive failure was F[5, 62] = 0·22. This was a good result. Using the same shortened sample, dynamic post-sample stability tests were performed as a final, but nonetheless important check against mis-specification. Dynamic forecasts were generated to 1988(3). No significant forecast errors were detected, and the Root Mean-Squared Prediction Error (0·0233) was low compared to the mean of the dependent variable (0·1051). Other measures of forecast adequacy were similarly satisfactory, which indicates that the findings of within-sample stability of (1.2) are matched by its ability to provide adequate predictions beyond the estimation period.

Equation (1.2) provides valuable information about wage inflation. Considerable inertia in real and money wage movements characterizes wage setting. Further, the equation is homogeneous with respect to prices and lagged wages, with past inflation rates affecting the current growth in wages slowly, reflecting the institutional lags introduced by the lengthy and retrospective NWC machinery. There is evidence that the recent (one-quarter) change in the inflation rate (CPI) has a positive impact on money wage inflation.

While the aggregate (labour market) hysteresis hypothesis does not receive statistical support (at the 5 per cent level), equation (1.2) confirms that the hysteresis effects based on capacity utilization are strongly defined.

The incomes policy variables are all significant and reveal that the different phases of wage guidelines exerted moderating but unequal influences

on wages growth. The statistical identification of the differential effects confirms the usefulness of this specification.

In choosing AWE as the dependent variable it was acknowledged that a more appropriate measure of unit labour cost would be average hourly earnings (AHE) which is the ratio of AWE to average hours worked per week AWH. Its use raises the possibility, however, that variation in the pressure variable might influence AHE, not directly through moderating wage demands but indirectly due to the inertia of AWE in response to quantity adjustment by firms (that is, variation in hours worked). Accordingly, we examined an added variable test on the annual change in average weekly hours $\Delta_4 AWH$. The insignificant t-statistic confirms the relative insensitivity of AWE to hours worked and hence confirms the predominance of quantity adjustments over price adjustments (see Okun, 1981).

The negative coefficient on the real wage term $\Delta_1 RW(-4)$ indicates that workers are content to moderate wage demands if they have had recent real wage gains, whereas they push for higher wages growth if real wage cuts have occurred.[8]

In summary, the wage setting is dominated by inertia, institutional resistance and hysteresis, with incomes policies acting as a restraining force. The aggregate labour market is not a significant influence on money wages.

Homogeneity of AWE with respect to Wages and Prices

Equation (1.3) is the tested-down version of (1.1) in real terms, so that the dependent and lagged dependent variables are four-quarter log changes of real wages.

$$\Delta_4 RW = +0.02 + 0.86\Delta_4 RW(-1) - 0.69\Delta_1 RW(-4)$$
$$(5.23) \quad (15.17) \qquad (7.24)$$
$$-0.33\Delta_2\Delta_4 CPI + 0.39\Delta_1 GUT - 0.02 IP1 - 0.02 IP2$$
$$(3.89) \qquad (2.91) \qquad (3.18) \qquad (3.85)$$
$$-0.02 IP3 - 0.01 IP4 - 0.05 IP5 - 0.02 IP6 + 0.02 IP7$$
$$(2.61) \quad (1.98) \quad (5.19) \quad (4.34) \quad (3.53)$$
$$+0.07 TD1$$
$$(5.68) \tag{1.3}$$
$$\bar{R}^2 = 0.89 \qquad \sigma = 0.012$$

[8]An alternative hypothesis is that the statistical significance of the real wage term is the outcome of mark-up pricing by firms rather than active real wage resistance by workers. Under a constant mark-up, real wages grow at the same rate as labour productivity and income shares are constant, irrespective of the time path of money wages. The significant variation of the wage share and the negative impact of most eras of incomes policy on the trend rate of real wage growth over the sample period undermines this hypothesis.

TABLE 2
EQUILIBRIUM GROWTH IN REAL WEEKLY EARNINGS
FROM EQUATION (1.3)

Constant Term and Intercept Dummies	Annual Equilibrium Growth of RW (per cent)
$C = 0.0169$	$+5.5$*
$IP1 = -0.0197$	-0.9
$IP2 = -0.0210$	-1.3
$IP3 = -0.0166$	$+0.1$
$IP4 = -0.0107$	$+2.0$
$IP5 = -0.0469$	-9.1
$IP6 = -0.0208$	-1.3
$IP7 = -0.0225$	-1.8

* decentralized adjustment.

Recent studies, including Coe and Gagliardi (1985) and Grubb *et al.* (1983), suggest that Australia does not suffer from any significant real wage inflexibility. Our results also suggest that significant real wage flexibility at the macroeconomic level does appear to exist in Australia. However, of importance is the source of this flexibility.

The studies noted above fail to distinguish between the different phases of "incomes policies" and erroneously include (given our findings) significant (aggregate) excess demand proxies in their specifications. Our results reveal that aggregate labour market demand proxies do not constrain real wage growth whereas the wage-fixing guidelines have had an important impact on the trend rate of real wage growth. In addition, our results suggest that the design of the institutionally-administered guidelines is also influential in restraining real wage growth. The respective magnitudes of the incomes policy variables estimated for the real wage per week equation (1.3) and the associated equilibrium rates of real wage growth, assuming constant steady-state growth rates of money wages and prices for each era, can be readily computed and are shown in Table 2.[9]

The second phase of the wage-fixing guidelines produced the greatest deflationary effect on real (and nominal) wage outcomes during the era 1975–1981. Indeed under the third and fourth eras in the early 1980s, equilibrium real wages exhibited positive growth and the fourth era heralded the collapse of the guidelines in 1981.

Despite the pressures which contributed to the re-emergence of collective bargaining, an examination of the residuals of all equations over this period and also an added variable test based on the most statistically significant catch-up variable CUP ($=1$, 1982(1); $=0$ otherwise) leads to the conclusion

[9] From equation (1.3) equilibrium real wage growth is the solution to $(1.0-0.8577)\Delta_4 RW^* + 0.6878\Delta_1 RW = C + IP_i$. This is readily solved when it is noted that $\Delta_4 RW^* = 4 \cdot \Delta_1 RW^*$.

(also made by Gregory, 1986, s.68) that there is no strong evidence of successful catch-up behaviour.

The recent period of incomes policy has had a far more profound effect on real wages. Incomes policies in this period ($IP5$, $IP6$ and $IP7$) have varied from the wages pause in 1983 ($IP5$), the period of delayed wage indexation with some terms of trade discounting ($IP6$), to the final two-tier guidelines where flat rates were awarded leaving a 4 per cent margin to productivity bargaining ($IP7$).

These results conflict with the findings of Phipps and Gregory with respect to the early phases of the guidelines. It should be emphasized that the econometric methods employed and diagnostic performance of our equation is superior to all the previous Australian studies.

An Increase in Capacity Utilization

Equation (1.2) indicates that measures to stimulate capacity utilization rates under the Accord Phase 2 ($IP7$) only exert moderate pressure on wage inflation. For example, if firms perceived a sustained increase in the demand for their output (perhaps as a result of expansionary fiscal policy) and increased capacity utilization by 2.5 per cent, then equation (1.2) would predict an instantaneous increase in the growth of average weekly earnings of about 1.10 per cent, *ceteris paribus* (for predetermined $\Delta_4 CPI$). The long-run impact of the higher rate of capacity utilization depends on the assumptions made about the magnitude of price inflation in subsequent periods, in response to the growth of AWE. It is easiest to assume that the "shock" to the growth of AWE through the increase in GUT occurs in the steady-state scenario of constant real wage growth, say of magnitude k. The direct impact of $\Delta_1 GUT$ is only temporary, however, so that if steady-state real wage growth is restored, it will be at the same level as before, *ceteris paribus*, but corresponding rates of price and money wage inflation may be higher. In both these steady states, therefore, the following relationship holds

$$\Delta_4 CPI = \Delta_4 AWE - k \tag{3}$$

Of importance, however, is the nature of price adjustment during the notional intervening phase of disequilibrium. To avoid arbitrariness and to identify the most inflationary outcome, the steady-state relationship (3) is imposed as the price adjustment equation.

It can be shown (see Appendix A) that, under homogeneity, substitution of (3) into a general, algebraic version of (1.2) yields a soluble difference equation for $\Delta_4 AWE$ for initial conditions which can be specified to incorporate the impact of the increase in GUT. The steady-state increase in the log of the four-quarter change in AWE, namely $\Delta\Delta_4 AWE^*$, can be written as:

$$\Delta\Delta_4 AWE^* = c_4 \Delta_1 GUT / (2 - c_1 - c_2) \tag{4}$$

where c_1, c_2, c_4 are the coefficients attached to $\Delta_4 AWE(-1)$, $\Delta_1 \Delta_4 CPI$ and $\Delta_1 GUT$ respectively in (1.2). The long-run increase in the four-quarter change in AWE for a 2·5 per cent rise in GUT, $\Delta\Delta_4 AWE^*$ lies within the following bounds, given the coefficient values in (1.2).

$$1·9\% < \Delta\Delta_4 AWE^* < 2·0\% \tag{5}$$

It can be shown, using simulation techniques, that lagged price adjustment brings about a smaller long-run increase in AWE, as would be expected. Thus, it can be argued that, under hysteresis with respect to capacity utilization, the wage adjustment process does not explode following an increase in capacity utilization.[10]

By contrast, in conventional wage equations, if the coefficients on the nominal variables on the right-hand side sum to unity, the long-run Phillips curve is vertical. Thus a rise in the pressure variable above its *natural level* causes an ever-increasing rate of inflation. The absence of a significant pressure variable in level form in our preferred specification denies the presence of such a Phillips curve. Indeed, over any particular wage fixation period, real wage growth is independent of the level of activity in the steady state.

V COINTEGRATION, ERROR CORRECTION and DYNAMIC MODELLING

Important developments have occurred in dynamic modelling in recent years concerning the existence and estimation of cointegrated vectors and their relation to error correction models (ECMs) (see, for example, Granger, 1981; Hendry *et al.*, 1986; Hall and Henry, 1987; and Engle and Granger, 1987). Related developments have also been made in the area of testing data for unit roots (see, for example, Dickey and Fuller, 1979, 1981; Evans and Savin, 1981, 1984; Nelson and Plosser, 1982; Sargan and Bhargava, 1983; and Phillips, 1987).

Mitchell (1989) compares a conventionally estimated ECM in real wages with a dynamic wage model using the two-step cointegration repre-

[10]A number of shortcomings can be identified with this approach. The endogeneity of price adjustment is inconsistent with the assumption of exogeneity in the estimation procedure. The calculation is based on the restoration of the steady-state growth of real wages, which is inconsistent with the notion of equilibrium as signifying the balance of forces with respect to income distribution. Equilibrium may not be restored after a shock to the system. Relatedly, different wage-fixing guidelines have been introduced over time to maintain the balance of forces but there is no guarantee that, for example, the two-tier system (*IP7*) will be maintained in the future. In short, the economic system is characterized by uncertainty. Two points can be made: (a) from a policy perspective, the impact of stimulatory policy on wage inflation is of key importance, particularly in the light of the Phillips curve explanation of inflation and (b) the parameter values from (1.2) reveal that at least 89 per cent of the change in four-quarter AWE growth occurs within the first four quarters so that the assumption of constant parameter values is not unrealistic.

sentation of the ECM. The results that bear on this paper are briefly noted below.

Testing for Unit Roots in Wage Equation Time Series

The properties of the individual time series should be established prior to testing for cointegration. The relevant variables in our model were tested for a unit root, which means that if $x_j = \rho x_{j-1} + \varepsilon_t$, stationarity requires that $|\rho| < 1$, whereas, if $\rho = 1$, the series is non-stationary (a random walk in this instance). The Augmented Dickey-Fuller (ADF) test uses a transform of the AR1 representation above to accommodate and test for a non-zero mean and drift (see Fuller, 1976, p. 373):

$$\Delta x_t = \alpha_0 + \alpha_1 t + \alpha_2 x_{t-1} + \sum_{j=1}^{k} \beta_j \Delta x_{t-j} + e_t$$

where $\Delta x_t = x_t - x_{t-1}$, $\alpha_2 = (\rho - 1)$, α_0 is a constant, and t is a linear time trend. k is often set at 2. This equation generates a test statistic (the t-statistic for α_2), which determines whether the level of x_t is stationary, that is, $x_t \sim I(0)$. Clearly, if $\rho = 1$, $\alpha_2 = 0$. If $\alpha_2 < 0$, then $\rho < 1$, and the process is stable and stationary.

If x_t is not integrated of zero order, then the ADF equation is transformed into second-differences, to test whether $x_t \sim I(1)$, that is, whether the first difference of x_t is stationary.

Table 3 reports the DF (being generated by imposing appropriate zero restrictions on the ADF equation), the ADF, and the Φ_2 and Φ_3 test statistics for the time series outlined above, in levels and first differences (see Dickey and Fuller, 1981). The variables are all quarterly and seasonally adjusted.

TABLE 3
TESTING FOR UNIT ROOTS IN WAGE VARIABLES

Wage Variable	X_0	T	DF	ADF	Φ_2	Φ_3
LAWE	60(2)	114	+8·23	−2·16	27·22	3·45
ΔLAWE	60(3)	113	−6·44	−5·76	11·08	16·62
LCPI	67(2)	86	−2·87	−2·68	45·31	5·34
ΔLCPI	67(3)	85	−5·19	−3·81	4·92	7·37
LGUT	60(2)	114	−1·61	−3·22	3·64	5·46
ΔLGUT	60(3)	113	−10·81	−6·13	12·54	18·80
LUR	60(2)	114	+0·76	−2·95	3·19	4·37
ΔLUR	60(3)	113	−7·24	−6·88	15·85	23·76
LAWH	66(2)	90	−0·54	−1·26	1·10	1·00
ΔLAWH	66(3)	89	−14·98	−7·71	19·83	29·75
LRW	67(2)	86	+2·43	−0·63	4·34	4·21
ΔLRW	67(3)	85	−8·12	−6·50	14·13	21·19

$\Delta = *_t - *_{t-1}$
$\Delta\Delta = (*_t - *_{t-1}) - (*_{t-1} - *_{t-2})$.

From Table 3, we can conclude that all the series are integrated of order 1, that is, $x_i \sim I(1)$. The fact that the first differences are stationary is a common property among economic time series. All the DF and ADF test statistics confirm the conclusion, which means that, in theory, the wage and price variables in level form could constitute a cointegrating set. We test for this in the next section.

The Φ_2 and Φ_3 test statistics also provide valuable information regarding the joint hypotheses relating to the process being a random walk with zero mean and no drift. Table V from Dickey and Fuller (1981) provides critical values for the test statistic Φ_2. The critical value for a sample size of 100 at the 5 per cent level is 4·88 (for n = 250, the critical $\Phi_2 = 4·75$). For the variables LGUT, LUR, LH and LRW, the joint hypothesis of a zero mean, no-drift random walk cannot be rejected. In first differences, the null can be rejected in each case.

Table VI, p. 1063 (Dickey and Fuller, 1981), provides critical values for the test statistic Φ_3. The critical value for a sample size of 100 at the 5 per cent level is 6·49 (for n = 250, the critical $\Phi_3 = 6·34$). The null of a driftless random walk cannot be rejected for all the variables. In first differences, the null is satisfactorily rejected in every case. From these tests it is fairly safe to say that all of the first-differenced series are stable and stationary autoregressive processes.

There is an interesting problem that these tests have raised with respect to their discriminatory power. It is well known that the tests have low power if $\rho = 0·9$ in a stable AR process. One of the variables tested was the capacity utilization variable LGUT, which is constructed from national accounts data, as the difference between actual GDP and a linked-peaks measure of potential GDP. By construction, GUT has an upper bound of unity (at the identified peaks). It is hard to imagine how the level of GUT could be anything but stationary, having no deterministic factor affecting its moments. However, the DF and ADF tests clearly show that LGUT is $I(1)$. This suggests that the test procedure has fairly low power for bounded time series.

Testing for Cointegration

Having established that all the variables in the wage model are of the same order of integration, that is $I(1)$, it is possible to test whether a linear combination of these individually non-stationary series can be found which is itself stationary. Table 4 reports the trials conducted. We should note that the DF and ADF tests are valid asymptotically and have low power in small samples. The critical values for the test statistics are from Engle and Granger (1987), and assume that T = 100, and the order of autoregression in the ADF equation is 4. The critical values (approximate as they are) for a 2-variable model: CRDW 0·386 (5 per cent) and 0·322 (10 per cent); DF −3·37 (5 per cent) and −3·03 (10 per cent); ADF −3·17 (5 per cent) and −2·84 (10 per

TABLE 4
COINTEGRATION TESTS
Sample: 1966(3)–1988(3)

| | Dependent Variable: LAWE | | | | | | Dependent Variable: LRW | | | | |
	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)	(viii)	(ix)	(x)	(xi)
CONSTANT	5·57	18·07	5·20	11·57	15·53	71·48	5·45	19·64	18·71	5·16	11·63
LP	1·23	1·03	1·02	1·01	1·06						
LGUT	0·51	1·50					−0·78	1·60			
LUR			0·20	0·11		−18·55					0·12
LH		−3·51		−1·75	−2·81			−3·95	−3·70	0·22	1·76
R^2	0·99	0·99	0·99	0·99	0·99	0·89	0·01	0·86	0·82	0·83	0·84
NOBS	89	89	89	89	89	89	89	89	89	89	89
CRDW	0·05	0·73	0·13	0·26	0·42	0·85	0·02	0·88	0·62	0·14	0·27
DF	−0·89	−4·42	−2·43	−2·99	−3·52	−4·75	−2·86	−4·98	−4·33	−2·52	−3·03
ADF	−1·08	−2·80	−2·83	−2·73	−2·64	−2·69	−2·31	−3·11	−3·25	−2·99	−2·78

cent). For a 3-variable model (and approximating > 3): CRDW 0·367 (5 per cent), 0·308 (10 per cent); ADF −3·13 (5 per cent), −2·82 (10 per cent).

As Hall and Henry (1988, p. 63) note: "It is clear that the critical values do not change enormously as we move from a 2-variable model to a 3-variable model. Nonetheless in the application given below we will be working outside the strict limits of these tables and so we must exert caution in the interpretation of these critical values."

Our ADF regressions were of a second order, except in the case of (vi) and (ix) which used a first-order AR process, due to the insignificance of higher order terms. In cases where the higher order terms are insignificant, the DF test is probably more reliable.

The results are generally inconclusive. Non-cointegration is accepted for equations (i), (iii), (iv), (vii), (x) and (xi). Equations (ii), (v) and (vi) reject the null using the CRDW and DF tests, although the ADF is marginal. Only (viii) and (ix) reject the null irrespective of the test. It is also true from (v) and (ix) that the activity variable LGUT is not required for cointegration. The fact that (viii) and (ii) cointegrate is probably due to the presence of LH rather than LGUT. The extremely tentative conclusion is that a subset of wage variables (LE, LP and LH) or (LRW and LH) may cointegrate, but no long-run real wage/activity relation is suggested.

The residual vector (Z) from each of the cointegrating regressions was then combined with differenced terms in an overparameterized model and a simplifying search procedure conducted. We note that a similar study by Hall and Henry (1986) wrongly used the current value of the cointegrating Z vector as their ECM, although they corrected this error in their 1988 work.

No satisfactory equation could be found with the two-step approach. The lagged (several were tried) Z vector (from each cointegration equation)

was only, at best, significant at the 10 per cent level. The diagnostics of the relevant equations were also inferior to the model presented in Section IV. The level variables in the various cointegrating equations were included in lagged form in the dynamic model but none was significant. These results indicate that there are problems with the cointegration tests.

Certainly, we are not prepared to cast aside the robustness of the model discussed in Section IV which used conventional methods to estimate the ECM representation in the face of our cointegration tests. We believe that the cointegration literature is of great conceptual importance and every effort should be made to include the tests in all dynamic models. However, as yet, the power and precision of the tests at the practical level is still low and thus doubt must exist when the results drawn from such work are to be interpreted. As Hall and Henry (1986, p. 236) suggest "...[none of]...these interpretations has so far been given a satisfactorily rigorous foundation, however, and so the two-stage procedure must be *ad hoc* at this stage" (emphasis added).

We prefer to be guided by the rule regarding the use of level variables, suggested by Hendry (1986, p. 204), who refers to "Hendry and Mizon (1978) who argued for retaining such variables in econometric models which were sufficiently well specified to have white-noise residuals, and hence avoid the spurious regression problem".

VI CONCLUSION

The results reveal that the different eras of wage-fixing guidelines can be statistically differentiated and are robust across different specifications. Except for the third and fourth phases of the guidelines ($IP3, IP4$) which signalled the end of centralized wage fixation in 1981, *incomes policy* successfully imposed a negative trend on the growth of real earnings, which was consistent with prevailing policy objectives.

There is no evidence of catch-up over the period after $IP4$ and prior to the wages pause. The behaviour of money and real wages over this period is consistent with the previous era of decentralized wage determination. Even over the period of the Accord, during which the unemployment rate fell from 10·1 per cent (1983(3)) to 7·0 per cent (1988(3)), real wages declined.

While our results provide evidence of real wage resistance and inertia in the growth of real weekly earnings in periods when the guidelines were not imposed, significant flexibility in macroeconomic wage outcomes is apparent when the guidelines are operating. Although this finding concurs with other studies (Grubb *et al.*, 1983), we would trace the source of flexibility to the guidelines themselves rather than appealing to conventional aggregate labour market pressure explanations.

The second important finding of this paper is that, in accord with Gregory, Dawkins and Wooden in Australia and Arestis, Wadhwani and

Henry and Ormerod in Britain, aggregate money wage outcomes are not determined by conventional market forces. The existence of a conventional Phillips curve relating inflation to unemployment is not supported. The annual growth of real weekly earnings is largely independent of conventional excess demand proxies and is strongly influenced by the prevailing institutional arrangements for wage fixing. The results support the work on wage determination which highlights the role of comparison and patterns of adjustment.

The pressure variable enters the simplified representation in a difference form which suggests that there is not a steady state or natural rate of unemployment. Indeed, the hysteresis property of the Australian wage equation is based on a firm-based measure of pressure rather than an aggregate labour market measure.

The change in capacity utilization contributes to the determination of aggregate wage inflation. This suggests that, because capacity utilization and profitability are highly correlated (using national accounts data reveals that the correlation between gross and net rates of return in all industries and GUT is approximately 0·79), the capacity to pay on the part of firms influences wage outcomes. Thus, following a recession, capacity utilization and profitability may improve as the firm restructures and higher wage outcomes may coexist with high unemployment (see Gregory, 1985, pp. 333–334).

The presence of hysteresis rather than a *natural* rate of capacity utilization suggests that, even under endogenous and instantaneous price adjustment, stimulatory policy will cause a higher but stable rate of money wage and price inflation rather than an ever-increasing rate which is associated with the expectations augmented Phillips curve. These tentative results point to the inappropriateness of counter inflationary policy based on deflationary monetary and fiscal policy. Conversely, from a steady-state configuration, an ongoing fall in inflation requires continuous declines in capacity utilization.

APPENDIX A

Under the homogeneity restriction, (1.2) can be written in the general form

$$\Delta_4 AWE = c_0 + IP_i + c_1 \Delta_4 AWE(-1) + (1-c_1)\Delta_4 CPI(-2) + c_2 \Delta_1 \Delta_4 CPI$$
$$+ c_3 \Delta_1 RW(-4) + c_4 \Delta_1 GUT \qquad (A.1)$$

Substituting for the inflation terms using (4), yields

$$\Delta_4 AWE = ((c_1 - c_2)\Delta_4 AWE(-1) + (1-c_1)\Delta_4 AWE(-2)$$
$$+ c_4 \Delta_1 GUT)/(1-c_2) \qquad (A.2)$$

since k satisfies $c_0 + IP_i - (1 - c_1 - c_3/4)k = 0$.

Deleting the term in $\Delta_1 GUT$ term yields a soluble difference equation in $\Delta_4 AWE$ which is readily solved for the initial conditions

$$\Delta_4 AWE(0) = \Delta_4 AWE^*$$

$$\Delta_4 AWE(1) = \Delta_4 AWE^* + (c_4/(1-c_2))\Delta_1 GUT \qquad (A.3)$$

as

$$\Delta_2 AWE(t) = \Delta_4 AWE^* + ((c_4 \Delta_i GUT/(2-c_1-c_2)))(1-[(c_1-1)/(1-c_2)]^t) \qquad (A.4)$$

Since $c_1 > c_2$ the long-run steady state increase in average weekly earnings growth following an increase in capacity utilization is

$$\Delta\Delta_4 AWE^* = c_4 \Delta_1 GUT/(2-c_1-c_2) \qquad (A.5)$$

DATA APPENDIX

All variables are obtained from the ABS NIF-10 Databank, September 1988 [1966(3)–1988(3)]. The variables employed and their respective codes, in brackets, are as follows:

AWE   average weekly earnings of non-farm wage and salary earners (*WAR$*).
CPI   consumer price index weighted average of 8 capital cities (*PCPI*).
U   total unemployment rate (*RNU*).
GUT   capacity utilization (*GUT*).
AWH   average weekly hours of wage and salary earners (*NHW*).
JVAC   total job vacancies (*NVA*).
ROT   overtime per employee (*ROT*).
LF   total civilian labour force (*NLF*).

In the estimated equations, all variables are expressed in logarithms and

$$\Delta j^* = {}^* - {}^*(-j)$$

$$\Delta ij^* = \Delta j^* - \Delta j^*(-i).$$

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Notes for Contributors

1. Articles should contain original unpublished work, not submitted for publication elsewhere.

2. Three copies of manuscripts should be submitted, with double-spaced type on one side of the paper only. They should be accompanied by a summary of the article of **not more than 100 words** which can be sent to the *Journal of Economic Literature* if the paper is accepted for publication.

3. Acknowledgements and references to grants, etc., should appear as a footnote to the author's name† and should not be included in the main list of footnotes. The author's name and institution should appear below the title.

4. Footnotes should be kept to a minimum and be listed consecutively throughout the text with superscript arabic numerals.

5. Diagrams should be in a form suitable for reproduction, with letters and symbols clearly printed.

6. Mathematical notation is very expensive to reproduce. In order to avoid unnecessary printing costs, the following guidelines should be noted.

    (i)   Avoid the use of unusual symbols where common ones are satisfactory.

    (ii)   Avoid fractions in the form $\frac{a}{b}$ and use instead a/b.

    (iii) Care should be taken to distinguish symbols which cannot be typed. Special care is required with letters and symbols such as w and ω, x and α, a and r and with the use of 0 and 1 where these can be read as either letters or numbers.

    (iv)   Avoid symbols above letters ($\bar{B}$) unless essential, and use acceptable alternatives (B^*) where possible.

    (v)   Where mathematical formulae are set out and numbered, these numbers should be placed against the right-hand margin as … (1).

    (vi)   Careful checking of the layout of all mathematical formulae, including alignment, centring, length of fraction lines and type, size, position and closure of brackets, etc., should be carried out before the final manuscript is submitted to avoid corrections at proof stage.

    (vii) Where it would assist referees, authors should provide supplementary mathematical notes on the derivation of equations.

7. References should be indicated by the author's name and date of publication, as Hicks (1965), and listed at the end of the article according to the style in the following examples:

    Hicks, J. R. (1965). *Capital and Growth*, London, Oxford University Press.

    Johnson, H. G. (1975). "The Monetary Approach to Balance-of-Payments Theory: A Diagrammatic Analysis", *The Manchester School*, Vol. 43, No. 3, pp. 220–274.

8. Authors are responsible for the careful checking of galley proofs. Corrections must be kept to an absolute minimum and the Editors reserve the right to charge for authors' alterations to proofs.

9. Manuscripts will not be returned unless requested by the author.