

ISSN 0819-2642  
ISBN 0 7340 1671 9



THE UNIVERSITY OF  
MELBOURNE

THE UNIVERSITY OF MELBOURNE  
DEPARTMENT OF ECONOMICS

RESEARCH PAPER NUMBER 730

JANUARY 2000

TRENDS AND CYCLES IN AUSTRALIAN STATE  
AND TERRITORY UNEMPLOYMENT RATES

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# Trends and Cycles in Australian State and Territory Unemployment Rates\*

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## ABSTRACT

*This paper has as its subject matter the behaviour of state unemployment rates over time. Arguments are presented which suggest that the common approach which entails regressing state or regional rates of unemployment on the national rate is not likely to yield much useful knowledge. As a positive contribution to the literature, this paper focuses on two things: first, the behaviour over time in the dispersion of state unemployment rates and their relationship with the business cycle and; second, tests for the presence of common trends and/or common cycles in the state unemployment rates. The results suggest that there is a case which can be made for regional policy in Australia.*

KEYWORDS: Regional unemployment Common trends Common cycles

JEL Codes: R230 1313 8241 9412

\* The authors are grateful to Guay Lim for helpful advice in relation to section IV of the paper.

## *I Introduction*

The main purpose of this paper is to examine the characteristics of state<sup>1</sup> unemployment rates in Australia. Their time paths of relative unemployment for the states exhibit a great deal of variety. Amongst other things, we look at the behaviour of the variance of the unemployment rate across states over time and we also perform tests to determine whether common trends or common cycles can be identified across the series. If common trends and/or cycles can be identified, it suggests that movements in unemployment across the various states reflect similar responses to common shocks. This would suggest there is little rationale for region specific policies. In contrast, the absence of common trends or cycles suggests that the states have faced shocks or have exhibited markedly different responses to them. In this event there may be a case for spatially differentiated unemployment policy.

We begin in section *II* with a discussion of the usefulness of what seems to have become the dominant approach to these matters, namely regressing regional unemployment rates on the national unemployment rate and using information gained as a basis for policy. In section *III* we consider similarities and differences between the economic structures of the states, using a measure of distance in economic space combined with cluster analysis. This is followed in section *IV* by a description of the behaviour of state unemployment relativities over time. We focus in particular on whether or not there is a relationship between a measure of the dispersion of state unemployment relativities and the business cycle. Having examined the underlying structure of the state unemployment rates, in sections *V*, *VI* and *VII* we consider whether common trends and cycles can be identified in the time-pattern of the

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<sup>1</sup> In this paper the term 'state' is used to refer to both states and territories.

unemployment series. Our analysis follows in the tradition of the expanding literature that seeks to identify the empirical characteristics of national and international business cycles (see for example, Blackburn and Ravn 1992; Fiorito and Kollintzas 1994; Backus and Kehoe, 1992; Backus, Kehoe and Kydland 1995; Christodoulakis, Dimelsis and Kollintzas 1995; Hess and Shin 1997). The contributions to this literature can be viewed as part of a general attempt to identify the processes that generate economic time series and degree of co-movement between key variables. The final section provides a summary of the results and presents our conclusions.

## *II Studying the Relationship between State and National Rates of Unemployment*

An important precursor to any theorising about state unemployment and regional policy is to understand clearly the time series properties of the unemployment rates themselves. One aim of this paper is to provide such 'benchmarking' information. At the same time, a basic starting point of the paper is that the regions are entities worthy of study in their own right and that regional unemployment issues can best be understood by examining directly the interactions between the regions. This approach is rather different to the typical research program in the area in Australia, and elsewhere, which presumes that regional unemployment issues can best be understood by examining the relationships between each of the regional rates and the national rate. Since we adopt a different approach in this paper, we begin by discussing the deficiencies in what seems to have become the standard approach.<sup>2</sup>

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<sup>2</sup> Groenewold and Hagger (1995), use a VAR model to look at the way in which unemployment growth rates (not unemployment rates) in different states interact over time. This is one of the few studies to look at relations between the states and not to simply look at connections in numerical terms between state unemployment and national unemployment.

The view that insight can be gained by (say) regressing regional unemployment rates (or any other variable) on its national counterpart was developed independently of each other by Thirlwall<sup>3</sup> (1966) and Brechling (1967) in the UK and the notion that regional unemployment is best seen in relation to national unemployment rates rather than other region's unemployment rates has since become widespread. The most recent paper<sup>4</sup> using this approach in Australia is that by Debelle and Vickery (1998).<sup>5</sup> After examining the relationship between regional and national unemployment rates<sup>6</sup> they report that "most of the movement in state unemployment rates can be explained by variation in the national rate. As evidence of this, the coefficient of determination between the state and national unemployment rates is generally high, varying between 0.75 and 0.90. That is, at least three-quarters of the variation in a state's unemployment rate is attributable to variations in national unemployment" (p 8). In presenting the conclusions of their research they write: "We [also] find that movements in the national unemployment rate explain most of the variation in state unemployment rates, suggesting that aggregate, rather than state specific factors, are most important in understanding Australia's high aggregate unemployment rate" (p 30).

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<sup>3</sup> See also Dixon and Thirlwall 1975, Ch 4.

<sup>4</sup> So far as we are aware the paper by Jeffrey and Webb (1972) was the first systematic attempt to apply the Brechling-Thirlwall approach to Australia. Groenewold (1991) is another recent paper in the Thirlwall tradition.

<sup>5</sup> In their recent survey paper covering the dimensions, structure and history of Australian unemployment, Borland and Kennedy (1998, p 89) are summarising Debelle and Vickery when they [i.e. Borland and Kennedy] remark that it is a "*feature*" of Australian unemployment that "state-level labour markets move quite closely together with the national market" (our emphasis). It is our contention that this is NOT in itself a 'feature' of the world (Australian unemployment). It is simply an artifact of the method of enquiry.

<sup>6</sup> This is actually only a minor part of their paper which is primarily concerned with the role of labour mobility in labour market adjustment and a tendency towards equalisation of regional unemployment rates.

size of (say) 100 we would not regard the correlation coefficient between two series as significant unless it exceeded 0.2. In the case we are considering, where the series are related via a weighting process, we have seen above that it would be necessary to add the weighting factor to the chance factor before reaching any judgement about the significance of a correlation. For example, with a weighting factor of 0.5 and a chance factor of 0.2, the correlation between the regional rate and the national rate would have to exceed 0.7 to be regarded as significantly different from zero.<sup>10</sup>

The problem discussed above applies directly in cases where the state unemployment rates are regressed on the national rate. In fact, this procedure becomes especially hazardous where the states are not of equal size. Again, the problems arise because the national rate is simply the weighted average of the state rates, where the weights are the proportion of the national labour force to be found in each state. These proportions vary considerably between the states. In 1988 the proportions of the national labour force present in each state was as follows:<sup>11</sup> NSW, 34.0%; VIC, 26.0%; QLD, 16.4%; SA, 8.5%; WA, 9.6%; TAS, 2.6%; NT, 0.9%; ACT, 1.8%. Given this, are we really learning anything when we find that unemployment rate for (say) NSW is more highly correlated with the national unemployment rate than is the unemployment rate for (say) Tasmania?

It is also important to recognise that when we regress state unemployment rates on the national rate, the results for the different states regressions cannot in fact be independent of each other as the coefficients, even when estimated in separate regressions, must be subject to an 'adding up restriction'. This can be seen with a

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region's share of the national labour force.

<sup>10</sup> Simulation exercises by the authors confirmed this point.

<sup>11</sup> The year 1988 is the middle year of the data period we use later in this paper.

simple example.<sup>12</sup> Suppose the nation is made up of only two regions. The national unemployment rate ( $N$ ) must be the weighted sum of the rates in the two regions  $R_1$  and  $R_2$  so that  $N = \alpha R_1 + (1 - \alpha)R_2$  (where  $\alpha$  is the proportion of the national workforce in region 1). If we then have a regression model (suppressing the constants and the error term) where  $R_1 = b_1 N$  and  $R_2 = b_2 N$ , it must be the case that  $N = \alpha b_1 N + (1 - \alpha)b_2 N$  from which it follows that  $b_2 = (1 - \alpha b_1)/(1 - \alpha)$ . Which is to say that  $b_1$  and  $b_2$  are subject to an adding-up restriction and so are not independent of each other or of their relative size of the regions (reflected in  $\alpha$ ). Suppose, for example, they are equal sized so that the coefficient  $\alpha = 1/2$ . In that case  $b_2 + b_1$  must sum<sup>13</sup> to 2 and the value of the two regression coefficients (even though they may be estimated in totally separate equations) cannot be independent of each other<sup>14</sup> as  $b_2$  must equal  $2 - b_1$ .

For the reasons given above, we are of the view that very little knowledge is to be gained by performing regressions of regional unemployment rates on national unemployment rates. Instead, we think it important to focus directly on the regional (State) economies and to inspect relationships between them in order to discern evidence of similarities or dissimilarities. However, before we do this it makes sense to devote some time to looking at the spatial and economic characteristics of the state economies. We do this for two reasons. First, to see if there is any reason to accept the presumption that they are so alike or interact so much that they can be regarded for policy purposes as one (national) entity. Second, so that later we can better judge

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<sup>12</sup> Johnston (1979) shows by empirical simulations that the slope coefficient is biased towards one the larger the region or town.

<sup>13</sup> This result that there is a connection between the sum of the regression coefficients and the number of regions (states) can be generalised to any number of regions.

<sup>14</sup> Brechling (1967) is aware of this. He notes that if there is one region for which the unemployment rate is less than the national rate there must be at least one other region for which the regional unemployment rate is greater than the national rate.

whether states with similar economic structures are experiencing similar (dissimilar) shocks and/or are responding to them in similar (dissimilar) ways.

### *III Differences (Distances) between the States in Economic Space*

The Australian states and territories may be regarded as regions of recent settlement for the purpose of understanding their economic history (see McCarty, 1965). As a result of this and also the timing of European settlement<sup>15</sup> in the various Australian colonies we find that Australia is one of the most urbanised countries in the world with the various state capital cities assuming increasing dominance over time. Another legacy of the past is that the Australian states are unusual as, with the exception of Tasmania, there are no second order cities in the state central place hierarchies (Logan and May, 1973). The result is that each of the state metropolitan areas interact strongly with their hinterland and also interact with each other (Burnley, 1980). Despite this, it is appropriate, as Stilwell (1974) has argued, to see each state as exhibiting its own urban hierarchy with metropolitan primacy.<sup>16</sup> It is also true that there are considerable physical distances between the state capital cities, a fact which itself tends to reinforce the strength of the interaction between state capitals and their hinterland but in the present context it is far more important to think of the distances between the states in economic space.<sup>17</sup> For example, we note that there are

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<sup>15</sup> The colonies were settled at the height of the development of nineteenth century capitalism. This above all else has had profound implications for the economic geography and the social structure in the states.

<sup>16</sup> The notion that there is relationship between the urban hierarchy and the business cycle has a very long history in regional economics. Key writers include Losch (1938), Vining (1946, 1949) and Isard (1960, Ch 6).

<sup>17</sup> The concept may be found in the brilliant but difficult paper by Perroux (1950). It refers to space in the manner in which a non-Euclidean geometer or mathematician would use the term. It is not geographic or physical space. For discussion of the concept see Perroux (1950) and Boudeville (1966) and the, by now, extensive literature on Growth Poles. For Australian readers we should point out that



differences in the export orientation of the states and also in their reliance upon public sector employment. The ratio<sup>18</sup> of International trade in goods - Exports to Gross State Product in 1987/88 was: NSW, 10.40%; VIC, 10.77%; QLD, 19.17%; SA, 10.16%; WA, 25.61%; TAS, 18.55%; NT, 28.13%, and; ACT, 0.03%. It is obvious that the export ratio is relatively high in QLD, WA, TAS and the NT. Estimates<sup>19</sup> of the percentage of employed wage and salary earners who are employed in the public sector June 1988 are: NSW, 27.2%; VIC, 27.6%; QLD, 30.1%; SA, 31.0%; WA, 30.7%; TAS, 35.1%; NT, 40.7%, and; ACT, 57.5%. To the extent that unemployment movements are related to demand shocks these figures alone suggest that some state economies are far more exposed to autonomous demand shocks than others.

If we look at the industry structure in more detail we see further differences between the states. Many industries are concentrated (localised might be a better word) in particular states. Tasmania, Western Australia, Queensland and South Australia have a relatively high proportion of their workforce engaged in Agriculture, forestry and fishing; Mining is disproportionately concentrated in the Northern Territory, Queensland, Western Australia and Tasmania.<sup>20</sup> There is very little Manufacturing in NT and ACT. Finance and insurance is relatively concentrated in NSW and Victoria. Tasmania is well under-represented in Property and business services. Government administration and defence is heavily concentrated in the NT and the ACT. Cultural and recreational services are over represented in QLD, the NT and the ACT. This suggests that the states are not 'close' to each other in economic

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the concept of economic space figures prominently in Geoffrey Blainey's famous 1966 book on Australian History titled *The Tyranny of Distance*.

<sup>18</sup> Source: ABS Time series in DX, Table 5220-7: Gross state product, expenditure components. The year 1987/8 is the middle year of the data period we use later in this paper.

<sup>19</sup> Source: ABS Labour Force data in DX.

<sup>20</sup> Mining is - not surprisingly - the most localised industry in Australia.

These two forces (dissimilar industrial structures and differing exposure to overseas influences on the one hand and high intermediate and aggregate demand linkages on the other) work in different directions and so it is impossible to make any *a priori* judgement as to the presence or absence of common trends and cycles across the various state economies. That also implies that we should not make any *a priori* judgment about the desirability of regional economic policy, including regional specific aggregate demand and/or labour market policy.

#### *IV State Unemployment Rate Relativities and the Business Cycle*

We observe a great deal of diversity in the levels and the time path of state unemployment rate relativities. The Figures set out below (2A-D) show the behaviour over time of the ratio of each states unemployment rate to the weighted average of all states (ie the national unemployment rate - for that is all it is, a weighted average of regional rates in this context). The states have been grouped into pairs according to the average level of their unemployment rate so that as we go from Figure 2A to 2D we progress from the pair of states with the lowest average unemployment rates (ACT and NT) to those with the highest (SA and TAS).<sup>28</sup>

[FIGURES 2A-2D NEAR HERE]

Clearly the time path or relative unemployment for the states exhibits great deal of variety. This is especially the case for Tasmania which is not only high on average but seems to be steadily trending upwards and WA which has a relatively high average but which seems to be trending downwards. It is also interesting to note

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<sup>27</sup> Groenewold and Hagger (1995) provide an interesting picture of those spillovers between states which may be construed as 'big-neighbour' effects.

<sup>28</sup> The mean unemployment rates (persons) over the period 1978-1999 are NSW 7.9%; VIC 7.7%; QLD 8.6%; SA 9.1%; WA 7.9%; TAS 9.5%; NT 6.7%; ACT 6.5%.

that the Dendrogram couplings (see Figure 1) are the same as the pairings according to unemployment rate from the pairs with the lowest unemployment rate to the pair with the highest. Taken by itself, this implies that states with similar industrial structures tended to have similar (mean) unemployment rates over the period.

Figure 3 shows the relative unemployment rate in all of the states shown in the one diagram. The advantage of doing this is that we might gain an impression of the presence or absence of compression or expansion of relative rates over time and/or over the business cycle.

[FIGURE 3 NEAR HERE]

There appears to be some indication of expansion of relative rates associated with recession episodes (81:2 - 83:2 and 89:4 - 93:3) and contraction of relative rates in between the two recessions and after 19993. Given this, it is useful to have recourse to statistical measures of dispersion so that we may better get a feel for movements over time and especially to consider the relationship between the dispersion of unemployment rates across regions and the stage of the business cycle.

The (weighted) Absolute Deviation<sup>29</sup> of state unemployment rates from the (weighted) average of the unemployment rates in all states taken together - that being the national rate,  $u_n$  - will be:

$$AD = \sum |(L_r / L)(u_r - u_n)|$$

Dividing this by the weighted average of the state unemployment rates gives a measure of Relative Dispersion (*RD*) (this is really a weighted Coefficient of Variation) which may be written as:

$$RD = \left[ \sum |(L_r / L)(u_r - u_n)| \right] [1 / u_n]$$

(UR) series.<sup>31</sup> Fitting an Error Correction Model with the dependent variable as the first difference in Relative Dispersion yields:<sup>32</sup>

$$D(RD) = -0.007 * D(UR(-1)) + 0.272 * D(RD(-1)) - 0.203 * (RD(-1)) - 0.172 + 0.010 * UR(-1)$$

(0.003)
(0.106)
(0.058)
(0.027)
(0.003)

Which is to say that there is a cointegrating relationship between the two variables such that in the long run an increase in the National Unemployment Rate by 1 percentage point tends to be associated with a decrease<sup>33</sup> in Relative Dispersion by 0.01. This would suggest that there is a trade off between dispersion (and thus 'equity') across states and low average (i.e. low national) unemployment. Our finding that there is a tendency for greater relative dispersion to be associated with (falling and) low unemployment periods is consistent with other Australian studies<sup>34</sup> including Groenewold<sup>35</sup> (1991), Stubbin & Hart<sup>36</sup> (1991), Andrews & Karmel<sup>37</sup> (1993), and Industry Commission<sup>38</sup> (1994).

<sup>31</sup> That is, the time series of the national unemployment rate.

<sup>32</sup> Where D is the first difference operator. The figures in parentheses under the estimated values of the coefficients are the estimated standard errors. Further tests using VEC show that it is appropriate to view the National Unemployment Rate as exogenous and the Relative Dispersion series as endogenous.

<sup>33</sup> This result is quite robust in that alternative lag lengths etc make very little difference to the long-run coefficient.

<sup>34</sup> The exception is in the study reported by Borland and Kennedy (1998). They examine the dispersion of unemployment in 107 DEETYA local labour markets in Victoria between 1984 and 1997. The simple correlation coefficient between the Coefficient of Variation and the Average Rate of Unemployment in the data reported in their Figure 11 (this covers 1984 through 1997) is (+) 0.14. (We are grateful to Jeff Borland for providing us with the data.)

<sup>35</sup> Looking at data from Feb 1978 through January 1990, Groenewold noted "the compression of the range (of state unemployment rates expressed as a ratio of the national rate) during the 1982/83 recession and the subsequent steady increase in dispersion as the national unemployment rate fell over most of the period since the recession [and up to January 1990]". Groenewold, (1991, p 17).

<sup>36</sup> Stubbin and Hart (1991, p 261f), look at the dispersion of the unemployment rate among 63 ABS Labour Force Regions over the period 1984 - 1990 (annual data). They report the Coefficient of Variation for each year. Comparing this with the average unemployment rate in each year we find that the simple correlation coefficient between the Coefficient of Variation and the average unemployment rate for their data set over the period is - 0.77.

<sup>37</sup> Andrews and Karmel (1993, p 48ff), look at the dispersion of the unemployment rate among 870 Local Government Authority and Statistical Local Areas over the period 1984 - 1991 (annual data). Since they report separately the average unemployment rate and the standard deviation it is possible to

The existence of a trade off between dispersion (and thus 'equity') across states and low average (ie low national) unemployment implies that as the national unemployment rate falls, micro and/or areally differentiated labour market policies<sup>39</sup> would need to bite harder (and effect proportionately more people) if equity is to be maintained.<sup>40</sup>

### *V Testing for Common Trends and Common Cycles*

The previous sections have examined the degree of similarity in the economic structure of the states, their differing unemployment experiences and the relationship over time between the dispersion of unemployment rates across the states and the level of unemployment in all of the states taken together. In this section and the two that follow we turn our attention to the time series properties of the unemployment data and the degree of co-movement exhibited by the states. In the literature cited in the introduction, the analysis of time series co-movement typically relies upon an examination of the contemporaneous and lagged cross-correlations between the variables, usually with initial pre-filtering of the data to ensure stationarity. Our approach is to examine the similarity between regional trends using the well-established tests for cointegration and the less well-known test for common cycles

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compute the Coefficient of Variation over time for their data set. The simple correlation coefficient between the Coefficient of Variation and the average unemployment rate over the period is - 0.87

<sup>38</sup> Industry Commission (1994, Vol 1, p 22 and p 125) report a Coefficient of Variation in regional unemployment rates on an annual basis over the period 1984 - 1992. They examine 61 sub-State regions - DEET small area labour markets - covering the whole of Australia. Their data suggests a negative relationship between the weighted average unemployment rate and the Coefficient of Variation over the period. The sample correlation coefficient between their two measures is - 0.84.

<sup>39</sup> Of course this might include policies related to mobility of labour. However, it must be noted that number of authors who have examined Australian data have reported that even after allowing for inter-state migration there remain permanent (or very persistent) differences in unemployment rates between the states. See Groenewold (1997) and Debelle & Vickery (1999) and the references cited therein.

<sup>40</sup> The relationship between RD and UR (and especially the possibility that the relationship has changed over time) is discussed at much greater length in Dixon, Shepherd and Thomson (1999).

introduced by Engle & Kozicki (1993) and Vahid & Engle (1993).<sup>41</sup> In comparison with the traditional correlation methods, the advantage of the test for common cycles is that it provides a stronger and more informative analysis of co-movement which emphasises the dynamic responses to innovations. In sections *VI* and *VII* we examine the properties of the state unemployment data and implement the tests for common trends and common cycles. Before doing so, we set out in this section relevant econometric theory and discuss the test procedures.

Most of the time series considered in macroeconomics exhibit non-stationary features and it is now common practice to implement tests for cointegration, to determine whether the variables in question follow common trends. The basic idea is that the presence of a cointegrating relationship provides evidence in favour of a long-run equilibrium relationship between the variables. The tests for common trends are generally based on the procedures introduced by Engle & Granger (1987), Stock & Watson (1998a) and Johansen (1988, 1991).

Importantly, from the point of view of this paper, Engle & Kozicki (1993) and Vahid & Engle (1993) have extended the notion of cointegration to a stationary setting, developing a test procedure that can be used to determine whether a range of stationary features are common across particular series that may or may not exhibit cointegration. As with the common trend analysis, the test for common feature represents an important advance in the identification of time series co-movement, extracting useful information from a detailed examination of statistical features traditionally regarded as a nuisance or a problem from the point of view of the classical regression model.

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<sup>41</sup> Previous contributions utilising this approach include Lippi & Reichlin (1994), Engle & Issler (1995)

The common trends and common features test procedures are closely related. This can most easily be seen by considering two non-stationary series  $Y_1(t)$  and  $Y_2(t)$  for which theory suggests there may be possible long-run and/or short-run relationships. We use the same symbol for both variables to emphasise that there is not necessarily any presumption about the direction of causality in the relationship. Following Stock and Watson (1988b) we regard the series as being generated by trend ( $T$ ), cyclical ( $c$ ) and noise ( $e$ ) components, such that:

$$Y_1(t) = T_1(t) + c_1(t) + e_1(t) \quad (1)$$

$$Y_2(t) = T_2(t) + c_2(t) + e_2(t) \quad (2)$$

If there is an observed degree of co-movement between the  $Y$  series, it must arise from one or more of the components ( $T$ ,  $c$  or  $e$ ) and the statistical questions that naturally follow are connected with the nature of the processes that generate these components and the extent to which they are similar over time. Concentrating first on the trend component, suppose that long-run movements in the series are generated by a common trend, where commonality is taken to imply proportionality. In this case, with the factor of proportionality represented by  $\alpha$ , the structure of the system is

$$Y_1(t) = T_1(t) + c_1(t) + e_1(t) \quad (3)$$

$$Y_2(t) = \alpha T_1(t) + c_2(t) + e_2(t) \quad (4)$$

and if the trend is common it should be possible to identify a linear combination of the original series  $Z(t) = Y_1(t) - \lambda Y_2(t)$  with no identifiable trend component. The linear combination is

$$Z(t) = T_1(t) - \lambda \alpha T_1(t) + c_1(t) - \lambda c_2(t) + e_1(t) - \lambda e_2(t) \quad (5)$$

and if  $\lambda = 1/\alpha$  this reduces to the stationary series:

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Caporale (1997) and Bai, Hall & Shepherd (1997).

$$Z(t) = c_1(t) - \lambda c_2(t) + e_1(t) - \lambda e_2(t) \quad (6)$$

The test for common trends is thus a test of whether there is some cointegrating vector  $[1, \lambda]$  for which the combined series (i.e. the  $Z$  series) is stationary.

To make the above procedure operational, it is necessary first to identify the precise form of the trend-generating processes in the two series. Since the trends can only be common if they are of the same statistical class, the first step in testing for cointegration is to determine whether they are. Assuming that trends of the same class have been identified, the second step is to determine whether or not they are proportionate. In practical applications in macroeconomics, the trend components are often stochastic trends and the appropriate steps involve testing first for the presence of unit roots in the individual series and then determining whether there is some linear combination of the variables that does not contain a unit root. Common or uncommon trends are then indicated according to whether or not a cointegrating vector  $[1, \lambda]$  can be identified for which the combined series (i.e. the  $Z$  series) is stationary.

If the  $Y_1$  and  $Y_2$  series are cointegrated, extraction of the common trend leaves a pair of stationary residual series generated by the remaining cyclical and noise components<sup>42</sup>  $c$  and  $e$ . If the series are not cointegrated, they do not share a common trend, but they can be rendered stationary by the application of differencing or some other filtering operation. Alternatively, if there are no trends present in the series, they are stationary at the outset and no pre-treatment of the data is required. Whichever of the above cases is applicable, we have a pair of stationary series and the interesting statistical questions are related to their underlying structure and the degree of co-movement between them.



Turning to the behaviour of the cyclical components,<sup>43</sup> the traditional approach adopted in the business cycle literature is to assess the degree of co-movement between the variables with reference to the contemporaneous and lagged cross-correlation coefficients. Engle & Kozicki (1993) and Vahid & Engle (1993) suggest an alternative means of determining the relationship between the cyclical components, based on an extension of the cointegration analysis to a stationary setting. The rationale behind the procedure can be seen if we look again at the  $Y$  series described by equations (1) and (2).

For the sake of argument, suppose the common trend has been extracted from the  $Y$  series or that they have been filtered in an appropriate manner so that we are left with the two stationary series:

$$y_1(t) = c_1(t) + e_1(t) \quad (7)$$

$$y_2(t) = c_2(t) + e_2(t) \quad (8)$$

One way to interpret the notion that a statistical feature is common across two series is to follow the example of the cointegration literature and assume that commonality is equivalent to proportionality. With the factor of proportionality denoted as  $\beta$ , this means that a cyclical feature is classified as common across the series if

$$c_2(t) = \beta c_1(t) \quad (9)$$

Forming a linear combination of the series,  $z(t) = y_1(t) - \lambda y_2(t)$  gives

$$z(t) = c_1(t) - \lambda \beta c_1(t) + e_1(t) - \lambda \beta e_2(t) \quad (10)$$

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<sup>42</sup> Strictly speaking the series are also driven by the disequilibrium error term implied by the cointegrating relationship.

<sup>43</sup> In this context a 'cyclical component' refers to a transitory but persistent process which might be common across several regions.

and if  $\lambda = 1/\beta$  the cyclical component is eliminated, leaving a series made up of the noise components. The procedure used to test for a common cycle is thus essentially the same as the one applied to the analysis of common trends, with the objective being to determine whether there is a common features vector  $[1, \lambda]$  that eliminates the stationary feature from a linear combination of the series. If such a common features vector can be identified, it provides strong evidence that (short run) movements in the series were driven by a common cyclical process.

Given that the common features analysis is essentially an extension of cointegration analysis, it is not surprising that the operational test procedures follow a similar line. In the test for cointegration, the first step is to determine whether the series are driven by a similar trend process, say an  $I(1)$  process, using the test for unit roots. In the common features setting, Engle and Kozicki (1993) suggest that the first step is to specify a statistical test  $s(y)$  to identify the stationary feature of interest and then to determine whether the feature is present in both the  $y_1$  and  $y_2$  series at some chosen significance level, such as 5%. Typical examples of the  $s(y)$  statistic might be the  $\chi^2$  LM and Wald-type tests used to identify features such as serial correlation and ARCH effects in the data, or the standard tests for features such as skewed distributions or excess kurtosis.

If the feature is identified in one of the series, but not the other, it indicates that they were driven by fundamentally different processes and that any further test for a common process is invalid. However, assuming that the feature has been identified in both series, the next thing to consider is whether a linear combination of the variables can be identified that does not exhibit the feature. The suggested procedure is to form a linear combination of the variables  $z(t) = y_1(t) - \lambda y_2(t)$  as

discussed above and then apply the feature test to the constructed  $y(t)$  variable for different values of  $\lambda$  until an estimate  $\hat{\lambda}$  is found that minimises the test statistic  $s(z)$ . This represents the estimate of the common features vector  $[1, \lambda]$  for the two series and a judgement about the presence or absence of a common feature can then be made according to how the estimated value of  $s(z)$  compares with the appropriate critical value. Engle and Kozicki (1993) demonstrate that the size of the test statistic applied at the second stage of the test is no greater than the size applicable to the two components of the constructed series at the first stage, implying that the same critical values can be applied at both stages of the test.

Since we are concerned with the cyclical behaviour of unemployment and the appropriate test statistic to consider is one that identifies the presence or absence of a cyclical feature. Vahid and Engle (1993, especially p 344ff) show that the presence of serial correlation in a stationary series implies the presence of a cyclical component and that the test for a common cycle in two series is essentially a test for the presence or absence of a common serial correlation feature. The procedure for two  $I(1)$  variables with no cointegration is to determine whether the first differences of the individual series exhibit serial correlation independently, and then to examine whether a linear combination can be identified for which it is absent. The first step involves testing for the presence of serial correlation in an AR or VAR model in first differences and is effectively a test of whether changes in the variables are predictable from their own history or from the joint information set.<sup>44</sup> In vector notation, the VAR model for the two stationary  $y$  series is

$$y(t) = \alpha + \beta y(t-1) + e(t) \quad (11)$$

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<sup>44</sup> The VAR model is more general since it allows for any system interdependence.

and the presence of serial correlation in both series is indicated if the *LM* test values derived from both equations of the VAR exceed the relevant critical value. In this case, the test values can easily be calculated from the VAR model as  $NR^2$ , where  $N$  is the sample size and  $R^2$  is the conventional coefficient of determination, with the *LM* statistic distributed as  $\chi^2$  with two degrees of freedom. For variables that exhibit cointegration, the procedure is the same as above, except that the preliminary VAR should also include an error correction term, in recognition of the fact that short-run changes are partly driven by the disequilibrium errors in the system.

If serial correlation is identified in both series, the next step is to test the null hypothesis that they are driven by a common serial correlation feature. This is done by constructing the linear combination  $z(t) = y_1(t) - \lambda y_2(t)$  and then deriving the estimate  $\hat{\lambda}$  that minimises the *LM* test statistic derived from an auxiliary autoregression on the combined series

$$z(t) = c + \delta z(t-1) \quad (12)$$

The null hypothesis of a common serial correlation feature is then accepted or rejected according to whether the *LM* test statistic is below or above the appropriate critical value, with the test statistic at the second stage distributed as  $\chi^2$  with one degree of freedom. Engle and Kozicki (1993) show that  $\hat{\lambda}$  can be derived from a two-stage least-squares (2SLS) regression of  $y_1(t)$  and  $y_2(t)$ , using lagged values of the variables as instruments. The absence or presence of a common cyclical feature is then indicated by whether or not serial correlation is present in the residual series from the 2SLS regression, judged according to the size of the standard *LM* statistic derived from an auxiliary autoregression. If there is no significant serial correlation in the residual series, it suggests that movements in the combined series are not predictable

from the past information set and that the component series were therefore generated by a common cyclical process. In contrast, the presence of serial correlation in the residual series indicates that the series do not share a common cyclical pattern.

### *VI Testing for Common Trends*

In this section we apply the tests for common trends to Australian State<sup>45</sup> unemployment data. The series in question are seasonally adjusted quarterly unemployment rates for persons covering the period 1978:2 to 1999:1 for the six states and the two territories of the Commonwealth.<sup>46</sup> Time series for the unemployment rate for each state and territory are graphed in Figure 5.

[FIGURE 5 NEAR HERE]

Following the tradition of the empirical business cycle literature, we begin with an overview of the correlation, persistence and volatility properties of the series.

Table 2 shows the contemporaneous cross-correlation coefficients for the unemployment rates of the various states. An approximate guide to the interpretation of the correlation coefficients is that they are significantly different from zero at the 5% level only if their absolute value exceeds  $2/\sqrt{N}$ , where  $N$  is the sample size. For the sample size considered here, this means that the correlations can be regarded as significant if they exceed 0.22.

[TABLES 2 AND 3 NEAR HERE]

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<sup>45</sup> As mentioned earlier, the word 'state' is used here to refer to 'states and territories'.

<sup>46</sup> The data was obtained from the ABS Time Series section of the DX database. The series is from 6203.0. Note: ABS does not publish seasonally adjusted data for NT and the ACT. The seasonally adjusted series for these two regions were computed using the Census X-11 option in EViews to seasonally adjust the original monthly data for these two regions. (For information on the procedures used by the ABS to seasonally adjust data see Appendix to the ABS publications 6203.0 for Feb 1996 and the 1983 issue of 1308.0.) Quarterly figures were arrived at by averaging the monthly figures.

The correlations shown in Table 2 indicate significant positive co-movement in the unemployment rates across all states, with exception of the ACT-NT and ACT-QLD relationships, which are insignificantly different from zero. The correlations are generally high across the states, with the exception of those between both ACT and NT and the other states, which are generally, lower than average. It may be noted that, with one exception, all of the pairings with the greatest similarity in their industrial structures identified with the aid of cluster analysis in section *III* show significant positive correlations in the contemporaneous behaviour of their levels of the unemployment rate. However, it must also be noted that pairings with the highest correlations are, almost without exception, not found between states with the greatest similarity in their industrial structures. All of which suggests that there is no simple relationship between industrial structure and the behaviour of levels of unemployment over time.

Additional information about the time series properties of the state unemployment rates is provided in Table 3, which shows traditional measures of persistence and volatility. The Table lists the first four parameters of the estimated autocorrelation function, which is usually taken to be a measure of the extent to which the series exhibit persistence to shocks. The Table also shows the standard deviation of each of the series, which gives an indication of their volatility. The AR parameters indicate a high degree of persistence in all of the series. There also appears to be a similar degree of volatility in each of the series, with the exceptions of VIC and ACT, which show respectively rather more and rather less volatility in unemployment than the other states.

The problem with interpreting the information in Table 2 is that the correlation coefficients are strictly meaningful only if the series are stationary. The AR parameters shown in Table 3 are suggestive of a unit root in most of the series, or at least a near-unit root process, and so it may well be the case that the positive correlations are largely spurious (Granger and Newbold, 1974; Phillips, 1986). The obvious way to test for stationarity is with the Augmented Dickey-Fuller (ADF) regression, which tests for the presence of a unit root in an autoregressive model of the series (Dickey and Fuller, 1979). Table 4 reports the ADF test statistics for each series. The test regressions are applied to the levels of the unemployment rates, with up to 4 lags in the regression model<sup>47</sup>. The starred result for each variable indicates the optimal lag length suggested by the Schwartz Information Criterion (SC). The critical values shown in the Table are derived from Mackinnon (1991).

[TABLE 4 NEAR HERE]

The results reported in Table 4 point to the presence of a unit root in most of the series, with the possible exceptions of NSW and WA. Having said this, many of the results are on the test borderline and it is possible that the series are stationary with a near-unit root or that they are fractionally integrated (DeJong and Whiteman, 1991; Kwiatkowski et al, 1992; Beran, 1994).

Although there may be some uncertainty about the order of integration of the series, previous studies of aggregate Australian unemployment suggest a unit root process (Mitchell, 1993; Crosby and Olekanski, 1998; Groenewold and Hagger, 1998; Gruen, Pagan and Thompson, 1999) and it is probably safest to proceed on the

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<sup>47</sup> Experimentation indicated no significant time trend factors in the data and the reported results are based on regression models that include only a constant. Essentially the same results were obtained from models that included both constant and time trend terms.

assumption<sup>48</sup> that this is true also at the state level, suggesting that the correlation results shown in Table 2 are potentially spurious. We note also that Groenewold & Hagger (1995, p 200) and Debelle & Vickery (1998, p 15 n8 and 1999, p 262, n3) find that all state unemployment rates are non-stationary. Taking this as our starting point, what we have to determine is whether common trends can be identified in the unemployment rates across the states.

We begin with a series of bivariate residual cointegration tests to determine whether common long-run trends can be identified across the various pairwise combinations of the states. If there are any significant long-run relationships in state unemployment movements, we should be able to identify at least some common trend elements. Following the procedure suggested by Engle and Granger (1987), we first estimated a set of bivariate regression models relating unemployment rates across pairs of states and then in each case tested for the presence of a unit root in the residuals using the critical values for residual cointegration tests suggested by Davidson and Mackinnon (1993). To allow for the possibility that the results may depend on the normalisation used in the cointegrating regression (Ng and Perron, 1997) we allowed the unemployment rate for each state to enter the regression first on the left-hand side and then on the right-hand side.<sup>49</sup> The results are given in Table 5.

[TABLE 5 NEAR HERE]

The results from the bivariate models show no strong evidence of any long-run cointegrating relationships between the series, although there are four marginal cases in which cointegration is suggested at the 5% level (for NSW-NT, NT-TAS, SA-VIC

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<sup>48</sup> Results presented by Smith and Yadev (1994) suggest that, for stationary fractionally integrated processes, a reasonable practical procedure is to treat the series as if they are  $I(1)$ .

<sup>49</sup> These tests do not imply anything about the direction of causality and so it is quite arbitrary which variable appears on the LHS of the regression.



and WA-SA). There is some minor variation in the test values according to the normalisation used in the cointegrating regression, but there are no cases in which the conclusion of no cointegration is reversed at the 1% significance level. However, there are four cases in which the change in specification reverses the no cointegration conclusion at the 5% significance level. These are the four previously mentioned cases for which cointegration is suggested at the 5% level and our interpretation of this mixed evidence is that the series are probably not cointegrated.

To make sure that the conclusion of no common trends does not depend on the bivariate structure of the model, we also tested for cointegration in a multivariate model, using both the Engle-Granger procedure and the VAR procedure suggested by Johansen (1988). In addition, to safeguard against any possibility that the results from the Engle-Granger procedure might depend on the choice of normalising variable, we again estimated a series of models, allowing the unemployment rate of each state to enter in turn as the dependent variable. Table 6 shows the results of the multivariate cointegration tests.

[TABLE 6 NEAR HERE]

The unit root tests indicate that the residuals are non-stationary in all cases, suggesting no cointegrating relationship. In the case of the Johansen procedure, following standard practice, we estimated a vector error correction model in first differences, with the appropriate lag structure determined by prior estimation of an unrestricted VAR model in the levels of the unemployment rates. The Schwartz criterion suggested that a one period lag in the error correction model was sufficient to capture the dynamics of the system. Based on a model of this order, the null of no cointegration is accepted or rejected according to the size of the likelihood ratio (LR)

statistic proposed by Johansen, using the critical values of Osterwald-Lenum (1992). The Johansen procedure is invariant with respect to the choice of the normalising variable and the LR test value indicates no cointegrating relationship, confirming the results of the Engle-Granger procedure.<sup>50</sup>

The various cointegration tests point to the conclusion that there are no common trends present in the levels of the state unemployment rates. Our conclusion is therefore that long-run movements in unemployment across the states have not followed a common trend path.

### *VII Testing for Common Cycles*

Given that the various unemployment rates are  $I(1)$  and are not cointegrated, the next thing to consider is whether any common cycles are present in the (stationary) first differences of the series. We begin with a brief look at the correlation, persistence and volatility properties of the first differences. These are shown in Tables 7 and 8. Since the first differences of the unemployment rates are stationary, the interpretation of the summary statistics is fairly straightforward.

[TABLES 7 AND 8 NEAR HERE]

The correlation coefficients for the first differences of the series indicate that there are a number of significant<sup>51</sup> positive correlations between changes in the unemployment rate across the states, with the notable exceptions of the NT, the ACT and TAS. There are no significant positive correlations between the NT and any other state. A similar picture holds for ACT and TAS, which show either insignificant correlations with other states or correlations that are only marginally significant.

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<sup>50</sup> For completeness we also considered higher order models. The conclusion of no cointegration was

Interestingly, only two of the pairings of states with the greatest similarity in their industrial structures show significant and positive correlations in the contemporaneous behaviour of their first differences in the unemployment rate. This supports the argument put by Groenewold that differing 'cyclical sensitivities' of state unemployment rates cannot be explained by reference (solely) to differences in their industrial structure (Groenewold, 1991).

Turning to the persistence and volatility measures in Table 8, the first-order AR parameter in particular indicates some cyclical persistence in the differences of the unemployment rates. As one would expect, the persistence is far less pronounced for the first differences than it was for the (non-stationary) levels. As far as the standard deviations are concerned, it appears that volatility is similar across the states, with the exception of NT, which shows rather less volatility than all the other.<sup>52</sup>

The correlations reported in Table 7 indicate that there is some contemporaneous relationship between unemployment rate changes across most states, and, since the data is stationary, we can take it that the relationships are meaningful rather than spurious. What the information does not tell us, however, is whether the first differences share the much stronger form of co-movement implied by common cyclical dynamics, meaning a common pattern of adjustment to cyclical shocks. To determine this we need to perform a test for common cycles.

The first step in the common cycles test is to examine whether cyclical dynamics are present in the data, in the form of serial correlation. Following the

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generally robust with respect to changes in specification

<sup>51</sup> Again, the benchmark value for the correlation coefficient is 0.22.

<sup>52</sup> An interesting point to note here is that the standard deviations are all significantly lower than they were for the levels of the unemployment rates. This reduction in volatility suggests that the operation of first-differencing the data is appropriate in the present context, since over-differencing would be more likely to generate an increase in the variance of the series (Mills, 1990).

procedures outlined in section V above, we begin by examining a series of AR(1) models, to determine whether serial correlation (and hence cycles) are present in the individual series.<sup>53</sup> For completeness and to allow for any system interdependence, we also test for the presence of serial correlation in a series of VAR(1) models. In both cases, the presence or absence of serial correlation is indicated according to whether the LM statistic is above or below the relevant 5% critical value. The results are reported in Tables 9 and 10.

[TABLES 9 AND 10 NEAR HERE]

The message from both Tables is that serial correlation is strongly present in all of the series, with the exceptions of TAS, ACT and the NT.<sup>54</sup> This implies that there are no significant cyclical dynamics present in the first differences of these series and that a test for common cycles in these cases is not therefore appropriate. In all of the other relationships, strong serial correlation features are indicated, suggesting a well-defined cyclical pattern in unemployment changes. We now consider whether these cycles are 'common'.

The first-stage results suggest that tests for common cycles are applicable across all of the states, except for TAS, ACT and NT. Following the procedure suggested by Engle and Kozicki (1993), we tested for the presence of serial correlation in bivariate linear combinations of the relevant variables, with the value of  $\lambda$  (the factor of proportionality) estimated with a 2SLS regression model, using lagged values of the variables as instruments. The presence (absence) of a common serial

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<sup>53</sup> For the seasonally adjusted data examined here the AR(1) model should be sufficient to capture the dynamics of the system. An examination of the Partial Autocorrelation Functions for the series indicated that they are best described by an AR(1) structure. The earlier cointegration tests also point to an AR(1) structure.

<sup>54</sup> For TAS the AR(1) model rejects the hypothesis of serial correlation at both the 1% and 5% levels, the VAR(1) model only rejects the hypothesis for all pairs at the 1% level.

correlation feature is indicated by whether serial correlation is absent (present) in an auxiliary autoregression of the residual series, judged according to the value of the *LM* statistic. To allow for the possibility that the results might be sensitive to the choice of normalising variable in the 2SLS regression, the bivariate models were estimated with the unemployment rate of each state entering in turn as the dependent variable. Table 11 shows the resulting *LM* values together with the sign of the  $\lambda$  coefficient generated by the 2SLS minimisation procedure, with pro or contra-cyclical behaviour indicated by a positive or negative value for  $\lambda$ . In this case, note that the choice of the normalising variable affects some of the test values, but there are no instances in which the conclusion is altered by the change of specification.

[TABLE 11 NEAR HERE]

Based on the 5% critical value of the *LM* test, the results suggest that common cycles are present in all of the relationships within the set of states encompassing NSW, QLD, SA, VIC and WA. Our assessment of the evidence is that unemployment changes follow a common dynamic cyclical pattern in the five largest states of the Commonwealth, NSW, QLD, SA, VIC and WA. For TAS, ACT and NT, there appear to be no common cyclical dynamics with the other states. This suggests that the largest states are responding in similar fashion to common shocks whereas TAS, NT and ACT are responding either in a dissimilar fashion to the same shocks<sup>55</sup> and/or to dissimilar shocks.

### *VIII Summary and Conclusions*

The typical research program in this area in Australia (and elsewhere) presumes that regional issues can best be understood by examining relationships between regional and national unemployment rates. For the reasons given in Section *II* we are of the view that very little knowledge is to be gained by performing regressions of this type, no matter how skilled the investigator. Instead, we think it important to focus directly on the regional (State) economies and to inspect relationships between them in order to discern evidence of similarities or dissimilarities.

The time paths of relative unemployment for the states exhibit a great deal of variety. Given this, we began by looking at the behaviour of the variance of the unemployment rate across states over time. Econometric tests suggest that (as is the case overseas) there is a cointegrating relationship between Relative Dispersion and the (weighted) Average Unemployment Rate across the states such that in the long run an increase in the National Unemployment Rate<sup>56</sup> tends to be associated with a decrease in Relative Dispersion. This implies that there is a trade off between dispersion (and thus 'equity') across states and low average (i.e. low national) unemployment. We then conducted tests for the presence of common trends and common cycles in state unemployment rates. With respect to common trends, the various cointegration tests pointed to the conclusion that there are no common trends present in the levels of the state unemployment rates and that long-run movements in unemployment across the states have not followed a common trend path. With respect to common cycles, our assessment of the evidence is that unemployment changes

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<sup>55</sup> But this is unlikely to be the case for all three as TAS has a markedly different industrial structure to

follow a common dynamic cyclical pattern in the five largest states of the Commonwealth, NSW, QLD, SA, VIC and WA. For TAS, ACT and NT, there appear to be no common cyclical dynamics with the other states.

The absence of common trends suggests that the long-run movements in unemployment have followed different paths across the states and that there may be a case to be made for region-specific policies designed to alter state equilibrium/structural unemployment rates. In contrast, the existence of common cycles for many of the states (in particular, the five largest), suggests that movements in unemployment across those states reflect similar responses to common cyclical shocks and that region-specific counter-cyclical policies (if such exist) are unnecessary for the large mainland states. However, the absence of common cycles for Tasmania and the two territories suggests that they have faced a different pattern of short-run shocks and/or have exhibited different responses to them. In particular in relation to Tasmania, which has an unemployment rate which is well above the national average, there does seem to be a case for regional specific counter-cyclical policies as well as policies designed to reduce the long term or equilibrium level of unemployment in that state.

The main conclusion from our study, based on our examination of both the relationship between dispersion and the business cycle and our tests for common trends and/or cycles, is that there is a case for targeted regional employment policy in Australia.

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the other two.

<sup>56</sup> The national unemployment rate is the weighted average of the State Unemployment Rates.

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TABLE 1

*(Pairwise) Coefficients of Regional Specialisation*

|     | NSW  | VIC  | QLD  | SA   | WA   | TAS  | NT   | ACT  |
|-----|------|------|------|------|------|------|------|------|
| NSW | 0.00 |      |      |      |      |      |      |      |
| VIC | 0.06 | 0.00 |      |      |      |      |      |      |
| QLD | 0.10 | 0.11 | 0.00 |      |      |      |      |      |
| SA  | 0.08 | 0.08 | 0.09 | 0.00 |      |      |      |      |
| WA  | 0.09 | 0.10 | 0.05 | 0.09 | 0.00 |      |      |      |
| TAS | 0.10 | 0.12 | 0.10 | 0.07 | 0.09 | 0.00 |      |      |
| NT  | 0.23 | 0.25 | 0.19 | 0.26 | 0.19 | 0.24 | 0.00 |      |
| ACT | 0.31 | 0.31 | 0.28 | 0.33 | 0.30 | 0.31 | 0.18 | 0.00 |

TABLE 2

*Contemporaneous Correlation Matrix for Unemployment Rates*

|     | NSW  | VIC  | QLD  | SA   | WA   | TAS  | NT    | ACT  |
|-----|------|------|------|------|------|------|-------|------|
| NSW | 1.00 |      |      |      |      |      |       |      |
| VIC | 0.79 | 1.00 |      |      |      |      |       |      |
| QLD | 0.90 | 0.75 | 1.00 |      |      |      |       |      |
| SA  | 0.87 | 0.92 | 0.82 | 1.00 |      |      |       |      |
| WA  | 0.83 | 0.73 | 0.76 | 0.76 | 1.00 |      |       |      |
| TAS | 0.79 | 0.80 | 0.78 | 0.83 | 0.61 | 1.00 |       |      |
| NT  | 0.58 | 0.38 | 0.59 | 0.42 | 0.61 | 0.46 | 1.00  |      |
| ACT | 0.29 | 0.59 | 0.21 | 0.51 | 0.27 | 0.37 | -0.12 | 1.00 |

TABLE 3

*Persistence and Volatility in Unemployment Rates*

|       | NSW  | VIC  | QLD  | SA   | WA   | TAS  | NT   | ACT  |
|-------|------|------|------|------|------|------|------|------|
| AR(1) | 0.95 | 0.97 | 0.94 | 0.93 | 0.94 | 0.92 | 0.78 | 0.86 |
| AR(2) | 0.87 | 0.91 | 0.85 | 0.83 | 0.85 | 0.85 | 0.62 | 0.73 |
| AR(3) | 0.75 | 0.83 | 0.73 | 0.72 | 0.72 | 0.78 | 0.51 | 0.61 |
| AR(4) | 0.62 | 0.74 | 0.61 | 0.60 | 0.57 | 0.69 | 0.39 | 0.54 |
|       |      |      |      |      |      |      |      |      |
| STD   | 1.66 | 2.10 | 1.47 | 1.32 | 1.39 | 1.76 | 1.63 | 1.09 |

AR(*k*) = *K*th-Order Autocorrelation Parameter. STD = Standard Deviation of Series

TABLE 4

*Stationarity Tests: Unemployment Rates*

|        | NSW    | VIC    | QLD    | SA     | WA     | TAS    | NT     | ACT    |
|--------|--------|--------|--------|--------|--------|--------|--------|--------|
| ADF(1) | -2.29  | -2.02* | -3.11  | -2.33* | -2.18  | -1.94* | -2.57* | -2.37  |
| ADF(2) | -3.05* | -2.38  | -2.20* | -2.44  | -3.15* | -2.06  | -2.27  | -2.69* |
| ADF(3) | -3.00  | -2.58  | -2.41  | -2.65  | -3.67  | -2.27  | -2.33  | -2.08  |
| ADF(4) | -2.63  | -2.27  | -2.40  | -2.58  | -3.06  | -2.56  | -1.71  | -2.02  |

ADF Test 1% and 5% Critical Values = -3.51 and -2.90.

Note: \* indicates lag length suggested by SC



TABLE 5

*Bivariate Residual Cointegration Tests*

|     | NSW   | VIC   | QLD   | SA    | WA    | TAS   | NT    | ACT   |
|-----|-------|-------|-------|-------|-------|-------|-------|-------|
| NSW | •     | -1.55 | -2.49 | -2.49 | -2.70 | -2.20 | -2.37 | -2.82 |
| VIC | -2.71 | •     | -2.36 | -3.35 | -1.64 | -2.13 | -1.54 | -3.11 |
| QLD | -2.54 | -2.37 | •     | -2.17 | -2.68 | -3.25 | -2.57 | -2.49 |
| SA  | -2.92 | -3.80 | -2.59 | •     | -2.49 | -3.22 | -2.45 | -2.46 |
| WA  | -1.41 | -1.35 | -1.35 | -3.35 | •     | -2.13 | -2.76 | -3.05 |
| TAS | -2.61 | -2.65 | -2.62 | -3.20 | -2.01 | •     | -2.55 | -2.08 |
| NT  | -3.58 | -3.07 | -3.70 | -3.27 | -3.81 | -3.42 | •     | -2.88 |
| ACT | -2.47 | -3.03 | -2.50 | -2.82 | -2.47 | -2.53 | -2.43 | •     |

Dependent Variable on Left-hand Side.

Residual Unit Root Test 1% and 5% Critical Values = -3.90 and -3.34

TABLE 6

*Multivariate Cointegration Tests*

| Normalising Variable | Engle-Granger Unit Root Test | Johansen Procedure LR Test |
|----------------------|------------------------------|----------------------------|
| NSW                  | -2.95                        | 150.81                     |
| VIC                  | -2.88                        | 150.81                     |
| QLD                  | -3.18                        | 150.81                     |
| SA                   | -4.62                        | 150.81                     |
| WA                   | -3.15                        | 150.81                     |
| TAS                  | -3.55                        | 150.81                     |
| NT                   | -4.42                        | 150.81                     |
| ACT                  | -4.35                        | 150.81                     |

Residual Unit Root Test 1% and 5% Critical Values = -5.25 and -4.71

Johansen LR Test 1% and 5% Critical Values = 177.20 and 165.58

TABLE 7

*Correlation Matrix for First Differences of Unemployment Rates*

|     | NSW  | VIC  | QLD  | SA   | WA   | TAS   | NT    | ACT  |
|-----|------|------|------|------|------|-------|-------|------|
| NSW | 1.00 |      |      |      |      |       |       |      |
| VIC | 0.63 | 1.00 |      |      |      |       |       |      |
| QLD | 0.63 | 0.60 | 1.00 |      |      |       |       |      |
| SA  | 0.56 | 0.59 | 0.54 | 1.00 |      |       |       |      |
| WA  | 0.56 | 0.57 | 0.48 | 0.36 | 1.00 |       |       |      |
| TAS | 0.30 | 0.33 | 0.27 | 0.19 | 0.25 | 1.00  |       |      |
| NT  | 0.24 | 0.18 | 0.10 | 0.02 | 0.18 | -0.14 | 1.00  |      |
| ACT | 0.30 | 0.16 | 0.08 | 0.19 | 0.25 | 0.21  | -0.14 | 1.00 |

TABLE 8

*Persistence and Volatility in First Differences of Unemployment Rates*

|       | NSW  | VIC  | QLD   | SA    | WA   | TAS   | NT    | ACT   |
|-------|------|------|-------|-------|------|-------|-------|-------|
| AR(1) | 0.48 | 0.54 | 0.46  | 0.35  | 0.42 | -0.03 | -0.17 | -0.04 |
| AR(2) | 0.42 | 0.40 | 0.22  | 0.17  | 0.44 | 0.01  | -0.07 | -0.04 |
| AR(3) | 0.22 | 0.31 | -0.01 | 0.08  | 0.32 | 0.09  | 0.01  | -0.19 |
| AR(4) | 0.07 | 0.10 | -0.15 | -0.08 | 0.09 | -0.14 | -0.22 | 0.01  |
| STD   | 0.46 | 0.47 | 0.42  | 0.43  | 0.42 | 0.58  | 1.04  | 0.57  |

AR( $k$ ) =  $k$ th-Order Autocorrelation Parameter. STD = Standard Deviation of Series

TABLE 9

*AR(1) Serial Correlation Tests: First Differences of Unemployment Rates*

|           | NSW   | VIC   | QLD   | SA    | WA    | TAS  | NT   | ACT  |
|-----------|-------|-------|-------|-------|-------|------|------|------|
| <i>LM</i> | 29.84 | 24.26 | 39.95 | 28.88 | 14.67 | 0.10 | 2.58 | 0.10 |

*LM* 1% and 5% Critical Values = 6.63 and 3.84 ( $\chi^2$  with one degree of freedom)

TABLE 10

*VAR(1) Serial Correlation Tests: First Differences of Unemployment Rates*

|     | NSW   | VIC   | QLD   | SA    | WA    | TAS   | NT    | ACT   |
|-----|-------|-------|-------|-------|-------|-------|-------|-------|
| NSW | •     | 26.74 | 30.17 | 24.42 | 21.44 | 22.64 | 24.63 | 20.29 |
| VIC | 24.84 | •     | 29.82 | 25.19 | 25.91 | 23.78 | 23.92 | 23.40 |
| QLD | 35.64 | 19.85 | •     | 35.17 | 20.37 | 35.08 | 35.11 | 38.05 |
| SA  | 17.64 | 19.03 | 30.12 | •     | 25.43 | 14.09 | 11.35 | 12.86 |
| WA  | 15.05 | 17.77 | 19.32 | 19.94 | •     | 14.46 | 16.90 | 16.55 |
| TAS | 6.44  | 6.13  | 1.11  | 0.25  | 6.34  | •     | 5.06  | 0.54  |
| NT  | 3.34  | 2.70  | 3.33  | 3.45  | 4.28  | 2.62  | •     | 3.36  |
| ACT | 0.31  | 4.37  | 0.21  | 5.31  | 0.62  | 4.61  | 1.72  | •     |

Dependent Variable on Left-hand Side.

*LM* 1% and 5% Critical Values = 9.21 and 5.99 ( $\chi^2$  with 2 degrees of freedom)

TABLE 11

*2SLS Common Cycle Tests for First Differences of Unemployment Rates*

|     | NSW         | VIC         | QLD         | SA          | WA          | TAS        | NT         | ACT        |
|-----|-------------|-------------|-------------|-------------|-------------|------------|------------|------------|
| NSW | •           | 0.23<br>(+) | 0.58<br>(+) | 1.42<br>(+) | 3.34<br>(+) | <i>n.a</i> | <i>n.a</i> | <i>n.a</i> |
| VIC | 0.22<br>(+) | •           | 3.57<br>(+) | 0.23<br>(+) | 1.02<br>(+) | <i>n.a</i> | <i>n.a</i> | <i>n.a</i> |
| QLD | 0.58<br>(+) | 3.57<br>(+) | •           | 0.03<br>(+) |             | <i>n.a</i> | <i>n.a</i> | <i>n.a</i> |
| SA  | 1.42<br>(+) | 0.23<br>(+) | 0.03<br>(+) | •           | 0.64<br>(+) | <i>n.a</i> | <i>n.a</i> | <i>n.a</i> |
| WA  | 3.14<br>(+) | 1.00<br>(+) | 2.44<br>(+) | 0.63<br>(+) | •           | <i>n.a</i> | <i>n.a</i> | <i>n.a</i> |
| TAS | <i>n.a</i>  | <i>n.a</i>  | <i>n.a</i>  | <i>n.a</i>  | <i>n.a</i>  | •          | <i>n.a</i> | <i>n.a</i> |
| NT  | <i>n.a</i>  | <i>n.a</i>  | <i>n.a</i>  | <i>n.a</i>  | <i>n.a</i>  | <i>n.a</i> | •          | <i>n.a</i> |
| ACT | <i>n.a</i>  | <i>n.a</i>  | <i>n.a</i>  | <i>n.a</i>  | <i>n.a</i>  | <i>n.a</i> | <i>n.a</i> | •          |

Dependent Variable on LHS.

Since there was no evidence of an AR(1) process in the time series for the first differences in the unemployment rate for TAS, NT and ACT (see Table 9 and the accompanying text), the common cycles test is not applicable to those states.

*LM* 1% and 5% Critical Values = 6.63 and 3.84 ( $\chi^2$  with one degree of freedom)

FIGURE 1

*Dendogram for Coefficients of Regional Specialisation*

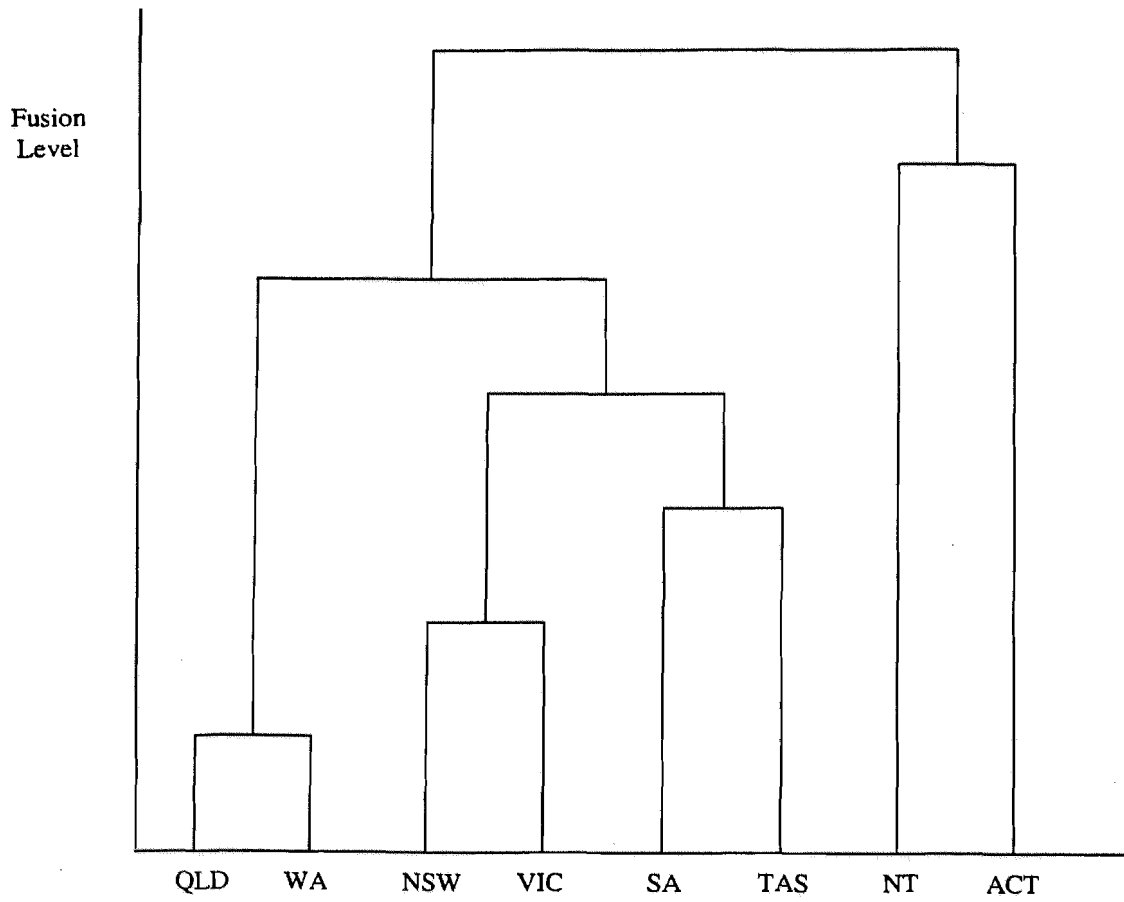


FIGURE 2A

*Time Path of Unemployment Relativities: ACT & NT*

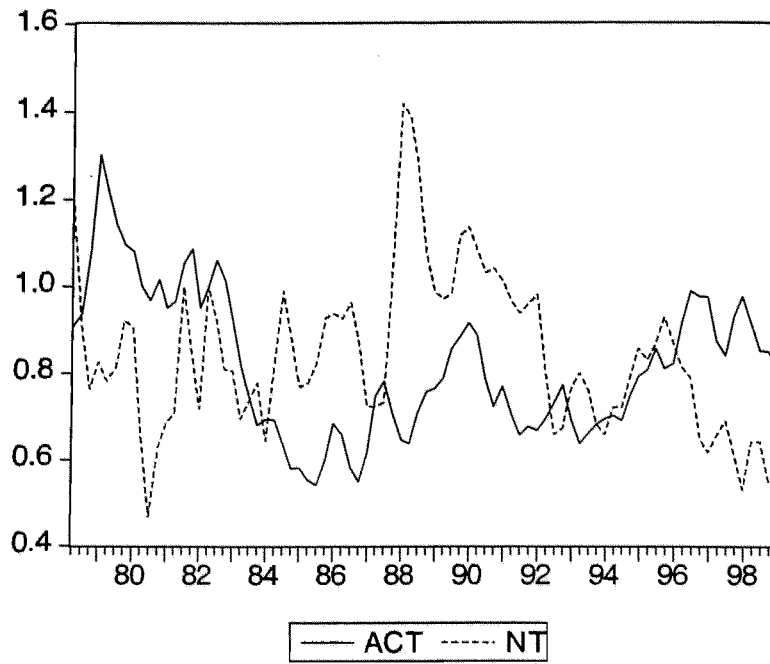


FIGURE 2B

*Time Path of Unemployment Relativities: NSW & VIC*

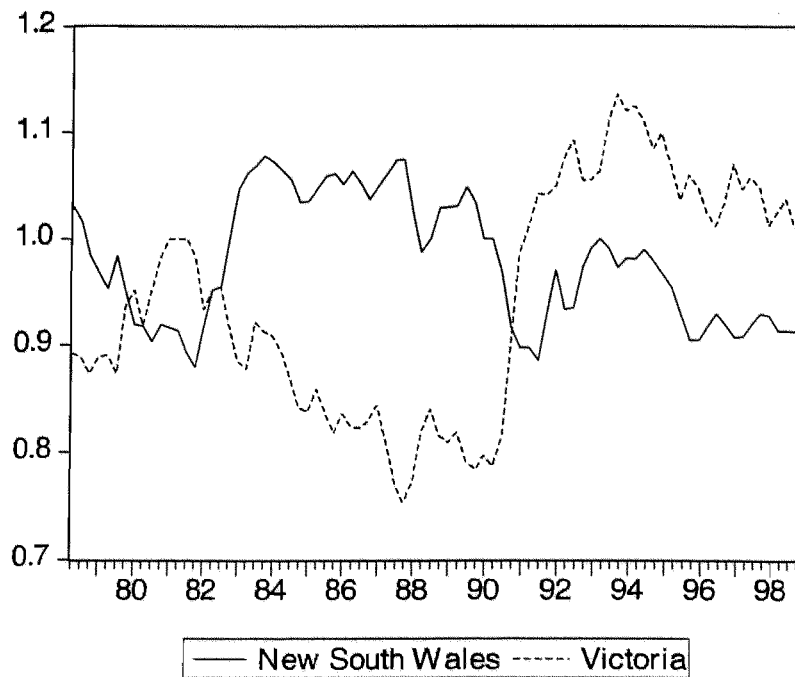


FIGURE 2C

*Time Path of Unemployment Relativities: QLD & WA*

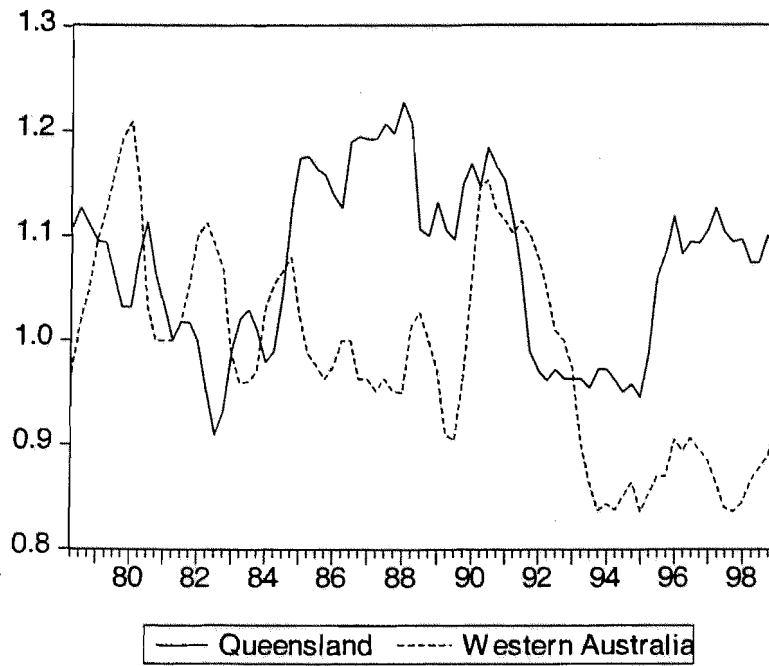


FIGURE 2D

*Time Path of Unemployment Relativities: SA & TAS*

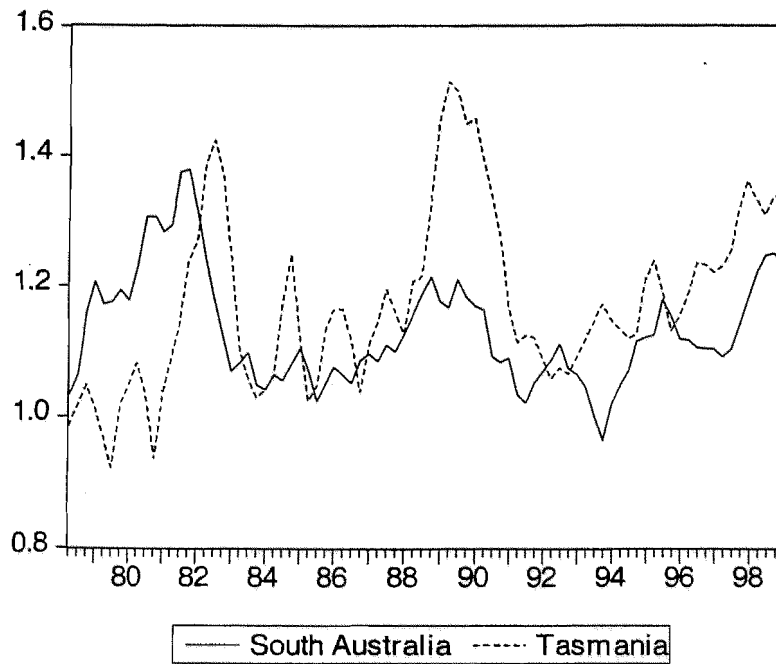




FIGURE 3

*Time Path of Unemployment Relativities: All States*

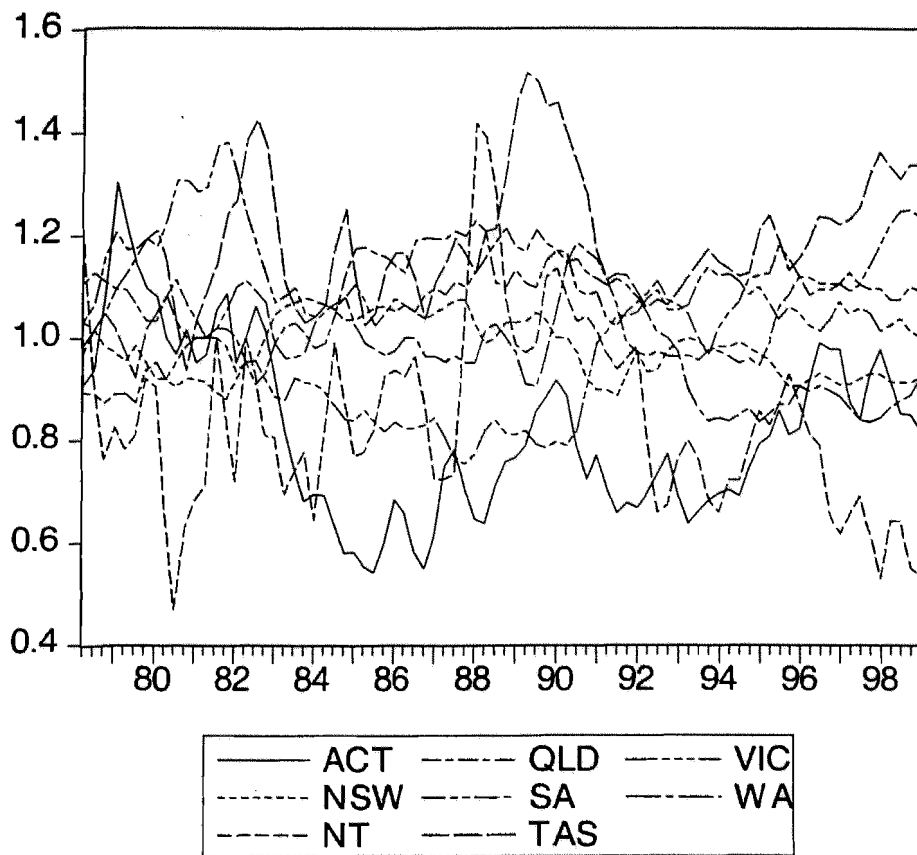


FIGURE 4

*Coefficient of Relative Dispersion in Unemployment Rates amongst the States (RD)*  
*and the Aggregate (Weighted Average) Unemployment Rate (UR)*

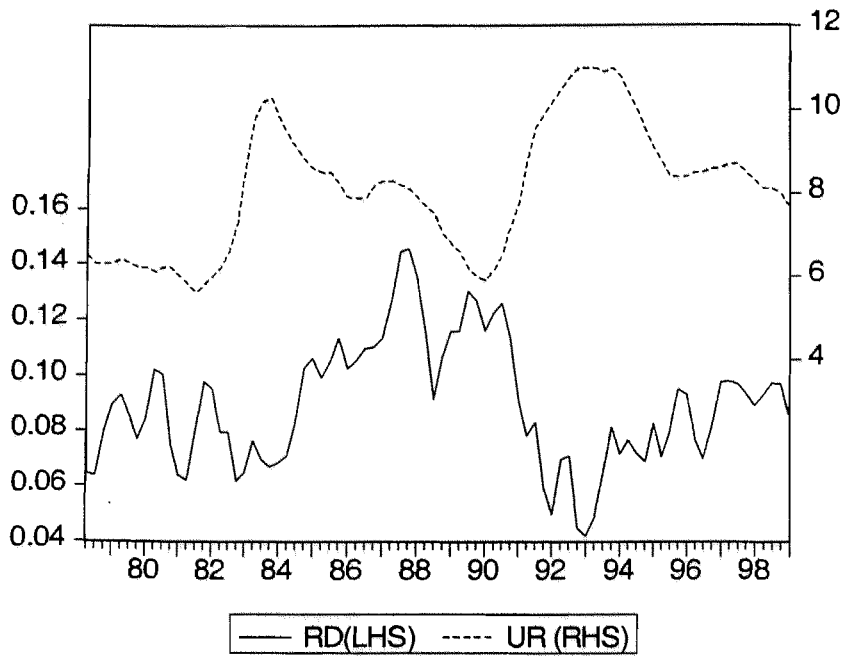
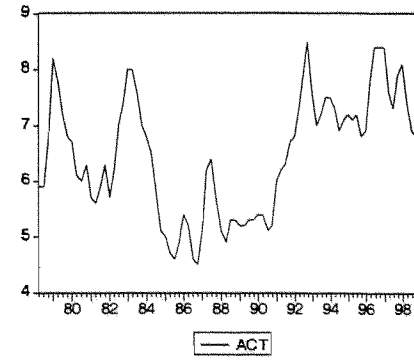
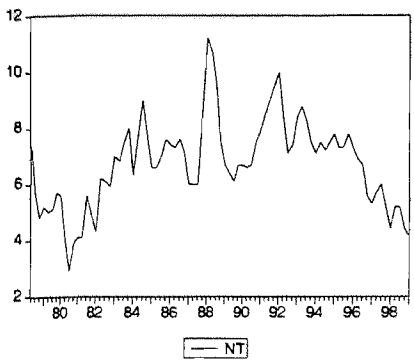
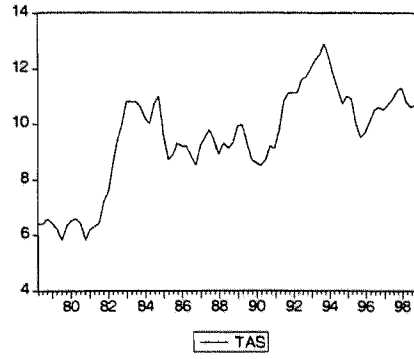
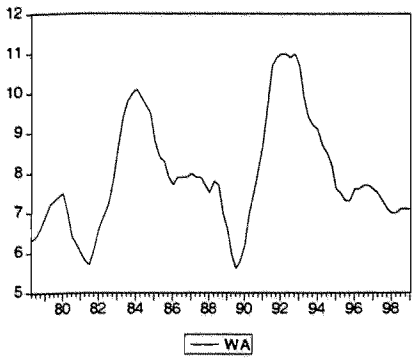
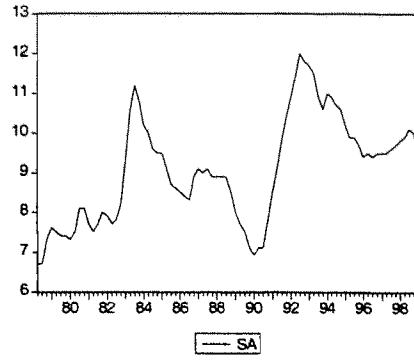
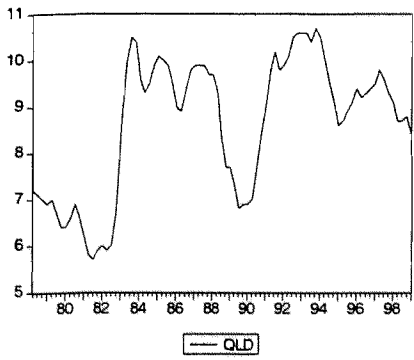
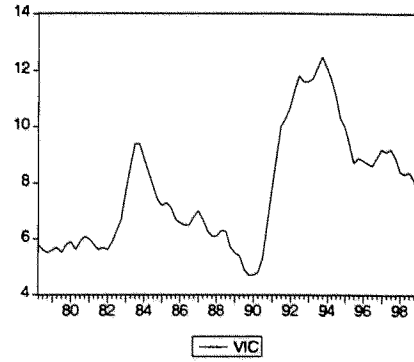
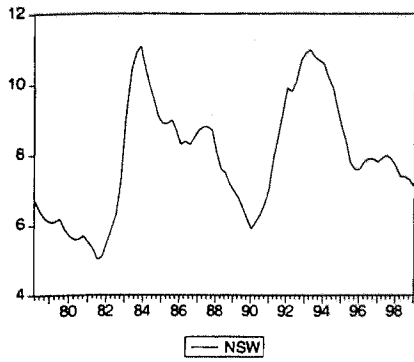


FIGURE 5

*Charts of State and Territory Unemployment Rates 1978:2 - 1999:1*



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