ISSN 0819-2642
ISBN 0 7340 1678 6

THE UNIVERSITY OF MELBOURNE

DEPARTMENT OF ECONOMICS

RESEARCH PAPER NUMBER 737

FEBRUARY 2000

REGIONAL UNEMPLOYMENT DISPARITIES

by

Robert Dixon
&
David Shepherd
&
James Thomson

Department of Economics
The University of Melbourne
Melbourne Victoria 3010
Australia.

Regional Unemployment Disparities

ROBERT DIXON*, DAVID SHEPHERD*# and JAMES THOMSON*

* Department of Economics, University of Melbourne, Victoria 3010, Australia
# Imperial College Management School, 53 Prince's Gate, Exhibition Road, London SW7 2PG United Kingdom

ABSTRACT

In this paper we examine the nature of disparities in regional (State) unemployment rates in Australia over the period 1978-1999 and their relationship to the national unemployment rate. As a measure of dispersion we use the sum of the (weighted) deviations of regional unemployment rates from the national rate. We show that this figure may be interpreted as the number of new jobs or labour force movements that would be needed to even out unemployment rates between regions, expressed as a proportion of the total number currently unemployed in all regions. Using co-integration analysis, we find that there is a (long-run) relationship between the degree of dispersion in the regional unemployment rates and the level of the national unemployment rate. The relationship between the two is negative implying that, as the national unemployment rate falls, micro and/or differentiated labour market policies need to bite harder (and affect proportionately more people) if equity in unemployment across regions is to be maintained. We also find that the trade-off between dispersion and unemployment has become steeper in the period following significant deregulation of the Australian economy in the early Nineteen-Eighties. It would appear likely that this reflects an increase in differences in the Natural Rate of Unemployment between the regions since that time.

KEY WORDS

Unemployment. Regional Disparities. Cointegration. Error Correction Model

INTRODUCTION

It has become common for researchers interested in regional unemployment to report (usually in the form of a graph) the relationship between a measure of the dispersion of unemployment rates across regions and the (weighted) average of those unemployment rates, which is the national unemployment rate (see MARTIN, 1997 and PEHKONEN & TERVO, 1998 for recent examples). However, other than reporting this information little seems to be done with it. In this paper we study the relationship between unemployment dispersion across regions and the national unemployment rate to see what it might tell us about the behaviour of regional (and national) labour markets. The series we examine are seasonally adjusted quarterly unemployment rates for persons covering the period 1978:Q2 to 1999:Q1 for the six states and the two territories of the Commonwealth of Australia.[1]

STATE UNEMPLOYMENT RATE DISPERSION AND THE BUSINESS CYCLE

There is a great deal of diversity in the levels and the time path of state[2] unemployment rates. Separate time series for the unemployment rate of each state and territory are graphed in Figure 1. Whilst Figure 2 shows the relative unemployment rate[3] in all of the states on one diagram. The usefulness of this plot is that it gives an indication of the degree of compression or expansion of relative rates over time and/or over the business cycle.

[FIGURE 1 NEAR HERE]

[FIGURE 2 NEAR HERE]

There appears to be some indication of contraction of relative rates associated with recession episodes (1981-83 and 1989-93) and expansion of relative rates in between these two episodes and after 1993. Given this, it is useful to have recourse to formal measures of dispersion, to get a more precise picture of movements over time and to facilitate a more considered analysis of the equilibrium relationship between the dispersion of unemployment rates across regions and the stage of the business cycle.

As our measure of dispersion we will use the Relative Dispersion (RD) of state unemployment rates around their weighted average in each period (i.e. the national unemployment rate). This may be written as:

$$\mathrm{RD} = \left[\sum (L_r / L_n) \left| u_r - u_n \right|\right] \left[1 / u_n\right] \tag{1}$$

where: L_r is the size of the labour force in region r.

L_n is the size of the labour force in all regions (the nation).

u_r is the unemployment rate in region r.

u_n is the unemployment rate in all regions (the nation).

An advantage in using the RD measure is that it has a very straightforward and intuitive policy-related interpretation. It is the number of persons in all regions taken together who would have to change their labour market status in order for all regions to have the (same) percentage unemployed as currently prevails in the nation, where that number (the total number whose labour market status would have to change) is expressed as a proportion of the total number currently unemployed in all regions. The easiest way to see this is to assume that there are only two regions (1, 2) and that they are of equal size, so that L_r/L is equal to 1/2 for both regions. In this event our expression for Relative Dispersion may be written as:

$$\mathrm{RD} = \left[\frac{1}{u_n}\right]\left[\frac{1}{2}\left|u_1 - u_n\right| + \frac{1}{2}\left|u_2 - u_n\right|\right] \tag{2}$$

Suppose that both regions have a (constant) labour force of 200, giving a national labour force of 400. Imagine that in region 1 there are 2 people unemployed and as a result the unemployment rate in region 1 is 1%. Suppose that there are 10 people unemployed in region 2 so that the unemployment rate in that region will be 5%. Given these figures the national unemployment rate will be 12/400 (= 1/2*1 + 1/2*5) which is 3%. If we calculate the value of RD for this data (using equation (2) above) we find that it is 2/3.

Imagine now that the labour market status of some individuals in both region 1 and region 2 changes so as to make the unemployment rate in both regions the same (i.e. 3%) while the national rate remains at 3%. Since 3% of 200 is 6 it must be the case that, in order for the unemployment rate in both regions to be 3%, an <u>extra</u> 4 people must become unemployed in region 1 and an <u>extra</u> 4 people must become employed in region 2. Notice that if we add together the number of people in both regions whose labour market status would have to change to equalise the unemployment rates at 3% we get the figure of 8 persons. If we divide this by the aggregate (national) number unemployed (3% of 400 = 12) we have 8/12 = 2/3, which is identical in value to the figure for RD arrived at above. All of which is to say that the measure presented in equation (1) of the Relative Dispersion of Regional Unemployment rates around their weighted average (the National unemployment rate) may be given a meaningful interpretation: the value of RD is equal to the number of new jobs or labour force movements that would be needed to even out unemployment rates between regions, expressed as a proportion of the total number currently unemployed in all regions.

STATISTICAL ANALYSIS AND RESULTS

Figure 3 shows how Relative Dispersion (RD) and the National unemployment rate (UR) have varied over time. Looking at the figure it would appear that Relative Dispersion is negatively correlated with the overall level of unemployment (and perhaps also the change in the level of unemployment).[4] The simple correlation coefficient between Relative Dispersion and the National Unemployment Rate (i.e. the weighted average of the unemployment rates in all of the states) over the period 1978:2 - 1999:1 is -0.45 and over the period 1984:1 - 1999:1 (this being the period after considerable deregulation occurred) is -0.83. However, we cannot use this finding as the basis for policy analysis without investigating the time-series properties of the data, since non-stationarity in the variables could well generate a spurious rather than a meaningful correlation between them (GRANGER and NEWBOLD, 1974; PHILLIPS, 1986).

[FIGURE 3 NEAR HERE]

An application of the Phillips-Perron and Dickey-Fuller tests indicated the presence of a unit root in both the Relative Dispersion (RD) and the National Unemployment Rate (UR) series. Since the variables are both I(1), we do indeed have the possibility of a spurious correlation between RD and UR and to determine whether this is so we need to examine whether there is a cointegrating relationship between the variables. If cointegration is found, it suggests that there is a meaningful long-run relationship between the variables and that they can legitimately be modeled in an error-correction form. If there is no cointegration, the correlation between the levels of the series is potentially spurious and the appropriate procedure is to examine the relationship in terms of a Vector autoregression (VAR) model applied to the first differences.

We begin by examining whether a cointegrating relationship can be identified for the two series over the full sample period. One way to do this would be to apply the procedure suggested by ENGLE and GRANGER (1987), which involves testing for the presence of a unit root in the residuals from an OLS regression of the two series. However, the more efficient procedure is the one suggested by JOHANSEN (1991 and 1995), based on full information maximum likelihood estimation of a Vector Error Correction Model of the series. In vector notation, the relationship between the levels of the series can be written as

$$Y_t = \Pi_1 Y_{t-1} + \ldots\ldots \Pi_k Y_{t-k} + e_t \tag{3}$$

where Y_t and e_t are $1\times N$ input vectors (in our case 1×2) and the Π terms are each $N\times N$ matrices of the system parameters. Equation (3) can be expressed in error correction form as

$$\Delta Y_t = \Gamma_1 \Delta Y_{t-1} + \ldots\ldots \Gamma_k \Delta Y_{t-k+1} - \Pi Y_{t-k} + e_t \tag{4}$$

where $\Gamma_i\ (i = 1....k)$ and Π represent respectively the parameter matrices on the first differences and levels of the series. In this framework, the presence or absence of a cointegrating relationship between the levels of the Y series can be assessed by examining the rank r of the Π matrix, using the likelihood ratio statistic suggested by Johansen, with cointegration indicated by $0 < r < N$. The results of applying the Johansen procedure[5] to the RD and UR series over the full sample are shown in Table 1.

[TABLE 1 NEAR HERE]

The results indicate that we are unable to reject the null hypothesis of no cointegrating relationship between the levels of the variables. At first sight this suggests that the correlation between RD and UR may be spurious and that the appropriate procedure is to examine the dynamic relationship between the (stationary) first differences of the variables, excluding any error correction term. Before proceeding down this path, however, it is instructive to look at Figure 4, which shows a scatter plot of the levels of RD and UR. The loops in the bottom LH corner all relate to the first part of the period which includes the years leading up to and including the recession of the early Nineteen-eighties. The line extending rightward from those loops ends near the RH vertical axis in 1983:Q4.

[FIGURE 4 NEAR HERE]

The scatter plot suggests that there may well have been an underlying structural shift in the relationship at the beginning of 1984 and that this may account for the lack of any observed cointegrating relationship over the full sample. As a test of this hypothesis, we re-estimated the model over two sample periods, from 1978:Q2-1983:Q4 and 1984:Q1-1999:Q1, to see whether a cointegrating relationship holds over these sub-samples. The choice of the break point for the sub-sample estimates may seem problematic since the usual test procedures, such as the Chow test, are based on the assumption that the data is stationary. In the present case, the series contain unit roots and hence the distributions of the relevant test statistics under the null are non-standard. There is a developing literature that deals with the detection of trend breaks in the context of univariate unit root testing (for example: PERRON, 1989; RAPPOPORT and REICHLIN, 1989; ZIVOT and ANDREWS, 1992; CHRISTIANO, 1992; BANERJEE, LUMSDAINE and STOCK, 1992; LUMSDAINE and PAPELL, 1997; BAI and PERRON, 1998). These studies are not directly applicable in the present context, in which we are testing the system for a cointegrating relationship, but they do suggest that the relevant critical values would be significantly higher than normal. They also suggest that relevant critical values are highly dependent on the precise structure and sample size of the model and can only be determined on a case by case basis, via extensive Monte Carlo simulations. Fortunately, for our purposes, the choice of an appropriate break point at around 1983:Q4/1984:Q1 seems quite clear on a number of grounds. First, the scatter plot of Figure 4 indicates a clear break in the relationship at the beginning of 1984. Secondly, the fact that the chosen break point allows us to establish a cointegrating relationship in itself provides evidence that the choice is appropriate. Finally, the choice of 1984:Q1 as the break point is also suggested by the significant changes in economic policy and institutional arrangements that occurred at the time in Australia (and elsewhere in the OECD). We come back to this last matter when we discuss our findings.

The results for the sub-sample estimates are shown on Table 2. They suggest that there is in fact a cointegrating relationship between RD and UR over both periods and that the relationship can therefore be examined legitimately in terms of a Vector Error Correction Model (VECM).

[TABLE 2 NEAR HERE]

The VECM and the estimated cointegrating equation for the RD-UR relationship over both periods are shown on Tables 3a and 3b, with the normalised cointegrating vector shown as *CE*. The t values associated with the parameter estimates are shown in parenthesis, with the associated probability levels in square brackets below. The models for both sample periods clearly passed the usual tests for serial correlation, functional form, normality and heteroscedasticity. Likelihood ratio tests also rejected attempts to restrict to zero any of the parameters of the cointegrating vector.

[TABLES 3a AND 3b NEAR HERE]

Comparing the cointegrating vectors for the two periods, we can see that the long-run relationship between RD and UR is negative. Our finding that there is a tendency for greater relative dispersion to be associated with (falling and) lower unemployment periods is consistent with other Australian studies including GROENEWOLD, 1991; STUBBIN & HART, 1991; ANDREWS & KARMEL, 1993 and INDUSTRY COMMISSION, 1994 and also with numerous studies for other countries. Interestingly and importantly from the policy viewpoint, our results also suggest that there has been a shift in the slope and position of the relationship, with a pronounced increase in slope over the later period and an associated increase in the intercept term.[6]

Turning to the full error-correction model, we can get some idea of the direction of long-run causality in the system by examining the magnitude and significance of the parameter on the lagged error correction term in both equations of the VECM. For example, if we found that the CE_{t-1} was significant in the ΔRD_t equation, but not in the ΔUR_t equation, it would suggest that adjustments to long-run equilibrium are brought about by movements in *RD* rather than *UR*, indicating a long-run direction of causality from *UR* to *RD* rather than *RD* to *UR*. For the 1978-1983 period, the CE_{t-1} term is significant at conventional probability levels only in the ΔRD_t equation, suggesting a long-run direction of causality from unemployment to dispersion. Having said this, the sample period for this equation is relatively short and the insignificance of the CE_{t-1} term in the ΔUR_t equation may in part be due to this. For the longer second period (1984-1999), the parameter on the error correction term is significant in both equations, and of similar magnitude, indicating that the direction of long-run causality runs both ways and that *RD* and *UR* should therefore be regarded as jointly determined in the system.

An alternative way to look at the causality issue is to consider whether there is any implied direction of causality in the short-run dynamics of the system. Concentrating on the second period, it appears from the t statistics that ΔRD_{t-1} is not significant in the equation explaining ΔUR_t but that ΔUR_{t-1} is marginally significant in the explanation of ΔRD_t. This provides some evidence that the direction of causation in the short-run dynamics may run from UR to RD rather than the reverse. A more formal way to test for this is to apply the standard Wald test, which allows us to determine whether valid restrictions can be placed on any of the parameters of the model. Tables 4a and 4b show the results for a range of parameter restrictions. The Wald statistic is distributed as χ^2 with k degrees of freedom, where k is the number of imposed restrictions. In this context, the Wald test is equivalent to a Granger-Causality test of whether the lagged variables on the right-hand side are significant in explaining movements in the left-hand side variable.

[TABLES 4a AND 4b NEAR HERE]

The parameter restriction tests confirm without exception the message conveyed by the t statistics for the equations reported in the previous Tables 3a and 3b. Concentrating on the longer second period (1984-99) we are unable to reject the restriction that ΔRD_{t-1} is not significant in explaining ΔUR_t. In contrast, the Wald test suggests that ΔUR_{t-1} does contribute to the explanation of ΔRD_t, although the test statistic is at the margin of conventionally-applied probability values. This means that there is some evidence of a short-run direction of causality running from unemployment to dispersion, in the sense that changes in unemployment generate subsequent changes in relative dispersion, but not the reverse.[7]

INTERPRETATION AND POLICY IMPLICATIONS OF THE RESULTS

We have seen that the long-run relationship between Relative Dispersion and the overall Unemployment Rate is negative. We need therefore to discuss three things. First, what it means for there to be a negative relationship between dispersion and the average unemployment rate. Secondly, the interpretation we can place on the slope coefficient which relates RD to UR and the policy implications of its magnitude. Thirdly, the economic interpretation of the intercept in the equation and its magnitude.[8]

We begin with the meaning of the negative relationship between RD and UR. Our explanation will be based on the following assumptions and/or stylised facts:

(a) We suggest that, as national unemployment rates become quite low, the value of RD primarily reflects differences in the Natural or Equilibrium rates of unemployment between the regions. (That there are differences in the 'equilibrium' or 'natural' rate across regions should not come as a surprise. A number of authors who have examined Australian unemployment 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.)

(b) As the economy moves into recession and the national unemployment rate rises above some low level, differences in the regional rates of unemployment reflect differences in *both* the natural rate and in the incidence of involuntary or demand deficient (cyclical) unemployment.

(c) Regional rates of unemployment are all positively correlated with the national rate, with the exact relationship between the two depending on the 'cyclical sensitivity' of the regional rates with respect to the natural rates (and the sensitivity of inter-regional migration to unemployment rate differences).

(d) The Absolute Dispersion of regional rates around the national rate (the Absolute Dispersion is $\left[\sum\left|(L_r / L)(u_r - u_n)\right|\right]$) is negatively correlated with the national rate of unemployment.[9]

(e) If Absolute Dispersion and the National Unemployment Rate are negatively correlated then it must be the case that Relative Dispersion and the National Unemployment Rate are negatively correlated.[10] So, in order to explain the negative correlation between Relative Dispersion and the National Unemployment Rate, we need (only) to explain the negative correlation between Absolute Dispersion and the National Unemployment Rate.

Given the above, to find that Absolute Dispersion (and thus Relative Dispersion) is negatively correlated with the National Unemployment Rate implies that regional differences in the Natural Rate of unemployment are greater (relative to the average level of unemployment) than are regional differences in the rate of Natural plus Involuntary unemployment (relative to the average level of unemployment). Again, given the assumptions and stylised facts presented above, the explanation must be that regions with relatively high Natural (structural) unemployment have relatively low coefficients of cyclical sensitivity whilst regions with relatively low Natural (structural) unemployment have relatively high coefficients of cyclical sensitivity.[11] If this is the case[12] then over some (relevant) range, dispersion must fall as the National unemployment rate (and all regions' unemployment rates) rise.

Having explained what we may infer from the existence of a negative relationship between Relative Dispersion and the National Unemployment Rate, we proceed now to talk about the magnitude of the slope coefficient of RD with respect to UR and its policy implications. We have already seen that the Relative Dispersion measure is equal to the number of new jobs or labour force movements that would be needed to even out unemployment rates between regions, where the number is expressed as a proportion of the total number currently unemployed in all regions. To find that the slope coefficient of RD on UR is negative says that a reduction in the national unemployment rate tends to be associated with an increase in the number of unemployed persons whose labour market status and/or region of residence would have to change if all states were (notionally) to have the same unemployment rate (of 5% in this example). This is an important and startling finding: it means that, as the national unemployment rate falls, the macro gain is offset by an increase in mis-allocation or inequality. The existence of this negative trade-off implies that micro and/or differentiated labour market policies would need to bite harder (and effect proportionately more people) during cyclical highs for 'equity' in unemployment to be maintained. However, we can go further. Our econometric results also suggest that the slope of the trade-off between Relative Dispersion and the (national) Unemployment Rate has increased over time, with a change in the slope parameter from –0.004 in the earlier period to –0.018 in the later period. The fact that the slope has increased between the first and second period implies that the loss of equity associated with any 1 percentage point fall in the national unemployment rate has risen markedly between the two periods. Indeed, if we calculate the elasticity of Relative Dispersion with respect to the National Unemployment Rate (at the means) for each of the periods, it is -0.348 for the period 1978-1983 and -1.656 for the period 1984-1999. This is a five-fold increase in the trade-off elasticity between the first and second period.

We turn now to the interpretation that may be placed on the magnitude of the intercept in the equation which relates Relative Dispersion to the National Unemployment Rate.

Although the (national) Unemployment Rate can never be zero, we suggest that the predicted value of RD when UR is at some relatively (arbitrarily) low number is informing us about (or is related to) the dispersion in the natural rates of unemployment in the regions.[13] Clearly, one cannot read off the absolute level of this from a single equation because we cannot be sure that all regional labour markets are operating at their natural rates at the same time or, if they are, at what National Rate this might occur. However, the relationship might be close enough that if the intercept were to change (say, rise) between two periods then we might be justified in concluding that the Natural Rates of Unemployment in the states had become more diverse. Our econometric results point to the fact that there has been an increase in the intercept term in the RD-UR relationship, from the beginning of 1984 onwards. The extent of the intercept shift is large, rising from an estimated value of 0.105 to 0.245. To the extent that the intercepts in the equations reflect disparities in the natural rates of unemployment between the regions, this suggests a marked increase in such differences.[14]

As mentioned earlier, it is of considerable interest that we find a structural break in the 1983-84 period. The Hawke Labor government was elected to Federal Parliament in early 1983. Shortly afterwards there was considerable financial deregulation, including the floating of the currency, removal of price support schemes for agricultural products, a marked reduction in tariffs on imports and an agreement entered into between the government and the Australian Council of Trade Unions (the peak union body in Australia) which led to marked wage restraint and (ironically) to greater centralisation and rigidity in wage-fixing arrangements. Whether intended to or not, these neoliberal economic rationalist polices "tended to intensify existing spatial inequalities" (TONTS, 1999, p 582). Also, along with many other countries, Australia experienced a marked change in Industrial Structure at this time with considerable employment losses in Textile Clothing and Footwear (the effect of the world-wide recession was exacerbated by large tariff cuts) and in Fabricated Metal Products and strong employment growth in various Service industries.[15] These changes had considerable region-specific impacts, especially in the case of the Textiles, Clothing and Footwear industries, which were largely concentrated in one state (Victoria).[16] It is all of these factors, taken together, which seem responsible for the structural breaks in the RD-UR relationship and for the worsening of the trade-off between Relative Dispersion and the overall Unemployment Rate which we have detected.

CONCLUSIONS

The Relative Dispersion of State unemployment rates is negatively correlated with the overall level of unemployment (and the change in the level of unemployment). This implies that there is a trade-off between dispersion (and thus 'equity') across states and low average (i.e. low national) unemployment and that the trade-off is one for one in terms of numbers of persons involved. In other words, as the national unemployment rate falls, micro and/or differentiated labour market policies would need to bite harder (and effect proportionately more people) for equity in unemployment across the regions to be maintained. We also find that the trade-off curve has become steeper in the period since significant deregulation of the economy in the early Nineteen-Eighties and that it is likely that differences in the Natural Rate of unemployment between the states has increased.[17] Taken together, these results suggest that there is a strong case for regional employment policy in Australia.

Acknowledgments - The authors are grateful to Jeff Borland, John Freebairn, Stephen Hall and Guay Lim for helpful comments.

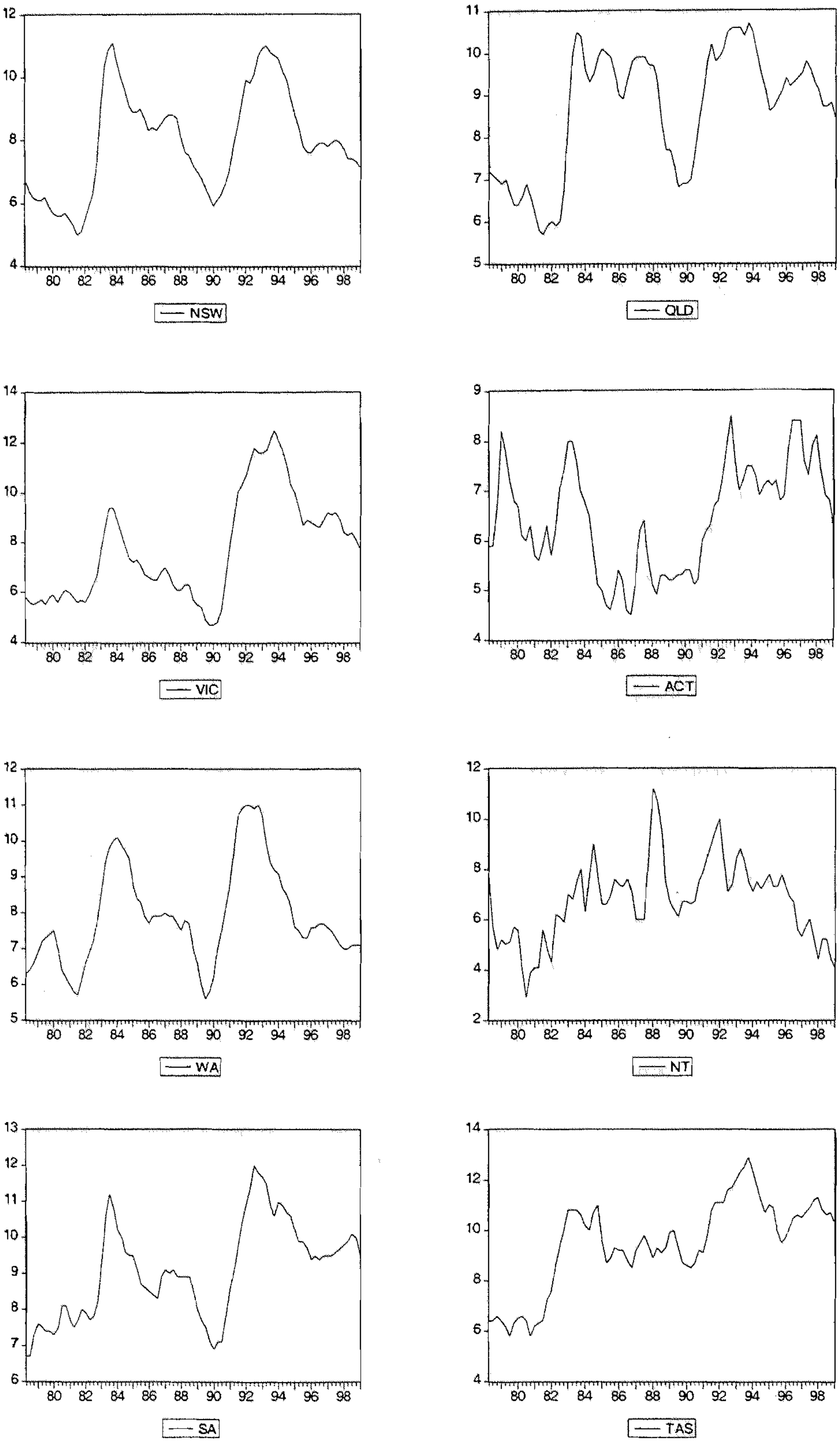

Fig. 1. Charts of State and Territory Unemployment Rates 1978:2 - 1999:1

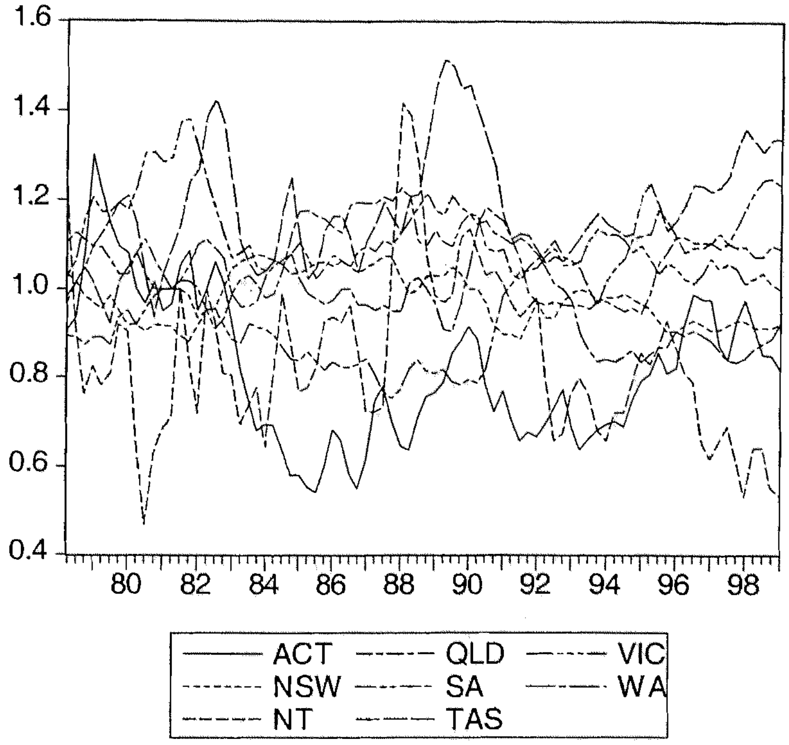

Fig. 2. Time Path of Unemployment Relativities: All States

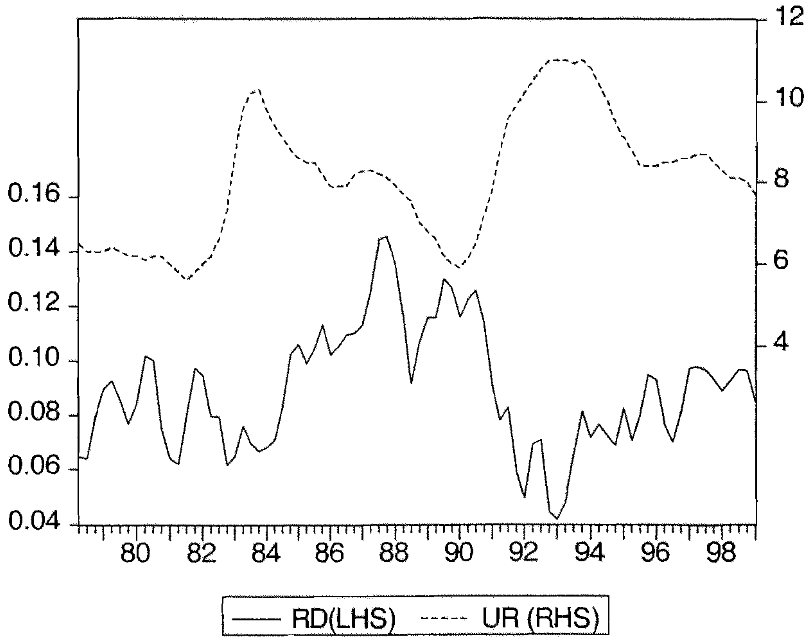

Fig. 3. Coefficient of the Relative Dispersion in Unemployment Rates amongst the States (RD) and the Aggregate (Weighted Average) Unemployment Rate (UR)

Table 3a. VECM for RD-UR: 1978-1983

Sample Period 1978:Q2-1983:Q4				
$\Delta RD_t = 0.667 \times \Delta RD_{t-1}$	$-0.003 \times \Delta UR_{t-1}$	$-0.418 \times CE_{t-1}$		$R^2 = 0.667$
(4.35)	(−0.80)	(−5.61)		
[0.000]	[0.432]	[0.000]		
$\Delta UR_t = 0.640 \times \Delta RD_{t-1}$	$+0.808 \times \Delta UR_{t-1}$	$-0.334 \times CE_{t-1}$		$R^2 = 0.615$
(0.11)	(6.12)	(−1.22)		
[0.911]	[0.000]	[0.239]		
$CE = RD_t + 0.004 \times UR_t - 0.105$				

Table 3b. VECM for RD-UR: 1984-1999

Sample Period 1984:Q1-1999:Q1				
$\Delta RD_t = 0.286 \times \Delta RD_{t-1}$	$-0.008 \times \Delta UR_{t-1}$	$-0.318 \times CE_{t-1}$		$R^2 = 0.221$
(2.13)	(−1.85)	(−3.00)		
[0.037]	[0.069]	[0.004]		
$\Delta UR_t = 3.219 \times \Delta RD_{t-1}$	$+0.844 \times \Delta UR_{t-1}$	$-0.475 \times CE_{t-1}$		$R^2 = 0.721$
(1.54)	(11.67)	(−2.86)		
[0.130]	[0.000]	[0.006]		
$CE = RD_t + 0.018 \times UR_t - 0.245$				

Table 4a. Wald Test of Parameter Restrictions: 1978-1983

Sample Period 1978:Q2-1983:Q4	
Test Regression: $\Delta RD_t = \alpha_1 \Delta RD_{t-1} - \alpha_2 \Delta UR_{t-1} - \alpha_3 CE_{t-1}$	
Parameter Restriction:	Wald Statistic [probability]
α_1=0	18.931 [0.000]
α_2=0	0.647 [0.064]
α_3=0	31.531 [0.000]
Test Regression: $\Delta UR_t = \beta_1 \Delta RD_{t-1} - \beta_2 \Delta UR_{t-1} - \beta_3 CE_{t-1}$	
Test Restriction:	Wald Statistic [probability]
β_1=0	0.013 [0.910]
β_2=0	37.424 [0.000]
β_3=0	1.484 [0.223]

Table 4b. Wald Test of Parameter Restrictions: 1984-1999

Sample Period 1984:Q1-1999:Q1	
Test Regression: $\Delta RD_t = \alpha_1 \Delta RD_{t-1} - \alpha_2 \Delta UR_{t-1} - \alpha_3 CE_{t-1}$	
Parameter Restriction:	Wald Statistic [probability]
α_1=0	4.566 [0.033]
α_2=0	3.435 [0.064]
α_3=0	9.015 [0.003]
Test Regression: $\Delta UR_t = \beta_1 \Delta RD_{t-1} - \beta_2 \Delta UR_{t-1} - \beta_3 CE_{t-1}$	
Parameter Restriction:	Wald Statistic [probability]
β_1=0	2.365 [0.124]
β_2=0	136.308 [0.000]
β_3=0	8.213 [0.004]

[1] The data was obtained from the ABS Time Series section of the Australian DX database. The series is from 6203.0. Note: ABS does not publish seasonally adjusted data for the Northern Territory and the Australian Capital Territory. 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 relevant monthly figures.

[2] In this paper the term 'State' is used to refer to both states and territories.

[3] By "relative" is meant the state unemployment rate in any period divided by the weighted average of all state rates (i.e. the national unemployment rate) for that period.

[4] This is a common finding in the literature. Indeed it is so common that MARTIN, 1993, p 798, refers to the "*customary* pro-cyclical movement in spatial unemployment differentials" (our emphasis).

[5] The lag structure of the VEC model was chosen by prior estimation of an unrestricted VAR model in the levels of the series, with the lag order selected by the AIC criterion. In this case, and later cases, a first-order lag in the VEC model was suggested.

[6] As an additional test of the structural break hypothesis, for the full sample we also estimated a model that included a slope dummy and a shift dummy. The estimated cointegrating vector for this model was: $CE = RD_t + 0.006UR_t + 0.01UR_t \times D_t - 0.126 - 0.103 \times D_t$, where D_t is a dummy variable that takes the value of zero before 1984Q1 and unity for the remainder of the sample. The parameters of this equation were all significant and the estimated parameter values are very much in line with those derived from the split sample estimates. In particular, the parameters on the unemployment terms show a negative relationship with relative dispersion, and the slope dummy suggests an increase in the slope of the relationship from –0.006 in the first period to –0.016 in the second. The model also suggests an increase in the intercept term, from 0.126 in the first period to 0.223 in the second. These results provide considerable support for the parameter values derived form the split sample estimates. There is, however, a note of caution to add, which is that, as we suggested earlier, the relevant critical values for use in the cointegration test are highly model-specific and the appropriate values that should be used for this particular model are not known and would have to be established from extensive simulation exercises. In view of this, we would prefer to emphasise the split sample estimates, for which the normal critical values are applicable.

[7] Having said this, the Wald test confirms that the CE_{t-1} term is highly significantly in both equations. This implies that RD_t and UR_t should ultimately be regarded as jointly determined in the system, and that the forces that generate long-run movements in unemployment also act to generate long-run movements in dispersion.

[8] Although, as we shall see, we can probably say more about changes in the intercept than we can about a particular level of the intercept.

[9] Over the whole period 1978:Q2 - 1999:Q1 the correlation coefficient between the two is +0.258. For the period 1978:Q2 - 1983:Q4 it is +0.682. For the period 1984:Q1 - 1999:Q1 it is -0.321. Since this suggest two different explanations of the relationship between RD and UR are required depending upon the time period, we have decided to focus on characteristics of the more recent period.

[10] Essentially this is because Relative Dispersion is simply the ratio of Absolute Dispersion to the National Unemployment rate.

[11] GROENEWOLD, 1991, p 26, summarising the results of his study of cyclical sensitivity using data for the states and territories over the period 1978:02 - 1990:10, writes: "high-unemployment regions tend to be less sensitive to the national cycle than low-unemployment regions". ELIAS, 1978, claims that this is also true for the UK, but c.f. GORDON, 1979. It is also important to note that there is no necessary reason why regions with relatively high natural rates of unemployment should have low cyclical sensitivity, however, looking at the estimates of BRECHLING, 1967; GROENEWOLD, 1991, and others, it does indeed appear to often be the case.

[12] Note that we have identified sufficient conditions for RD and UR to be inversely related. Other conditions would also suffice to give this result but we are concerned to present an explanation which is consistent with the Absolute Deviation also being inversely related to UR, because that appears to have been the case in Australia over the past 15 years.

[13] As mentioned earlier, a 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.

[14] FAHRER and PEASE, 1993, p52, find a significant outward shift in the Beveridge Curve for Australia in 1983/84.

[15] These changes (and their regional implications) are well documented in PRODUCTIVITY COMMISSION, 1998.

[16] In 1983, 58% of all employment in TC&F was to be found in Victoria yet the state had only 27% of the national labour force.

[17] One further point can be made. As ARCHIBALD, 1969; LIPSEY, 1960 and THIRLWALL, 1969, amongst others, have pointed out one would expect the aggregate Phillips Curve to shift in or out depending upon the relative levels of excess demand in the 'micro' markets. To the extent that our measure of Relative Dispersion is informing us about this matter one would expect the economy-wide inflation unemployment trade off to have shifted outwards between the first and second of the two periods we have examined. All of which is to say that it is in the widening of the differences in the regional natural rates that the increase in the economy-wide natural rate commented on by so many authors (for a recent survey see BORLAND and KENNEDY, 1998) may be seen to have its origins. We note that CROSBY and OLEKALNS, 1998, report an outward shift in the long-run Phillips curve between 1983 and 1984.

RESEARCH PAPER SERIES - RECENT PUBLICATIONS IN THE DEPARTMENT OF ECONOMICS

NO.	AUTHOR/S	TITLE	DATE	INTERNAT. WORKING PAPER NO.	ISBN NO.	TOTAL NO. OF PAGES
732	J. G. Hirschberg & P. J. Lloyd	An Application of Post-Dea Bootstrap Regression Analysis to the Spill over of the Technology of Foreign-Invested Enterprises in China	January 2000	IWP 669	0 7340 1673 5	29
733	Chris Brooks & Olan T. Henry	The Impact of News on Measures of Undiversifiable Risk: Evidence from the UK Stock Market	February 2000	IWP 670	0 7340 1674 3	28
734	Joshua S. Gans & Stephen P. King	Mobile Network Competition, Customer Ignorance and Fixed-to-Mobile Call Prices	February 2000	IWP 671	0 7340 1675 1	31
735	Jose-Luis Negrin & Charles E. Hyde	Access Pricing: A Survey	February 2000	IWP 672	0 7340 1676 X	35
736	Rob Feeney & Stephen P. King	Sequential Parimutuel Games	February 2000	IWP 673	0 7340 1677 8	12
737	Robert Dixon, David Shepherd & James Thomson	Regional Unemployment Disparities	February 2000	IWP 674	0 7340 1678 6	26