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Centre of Full Employment and Equity

Working Paper No. 01-04

Measuring persistence in unemployment rates

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May 2001

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1. Introduction

In this paper, we explore the notion of persistence with specific reference to unemployment rates. The persistent of shocks to real time series is now the subject of many studies. Nelson and Plosser (1992) provided evidence to support the notion that shocks to time series were probably permanent. They also argued that the cycles in real output, for example, were largely explained by these shocks. Mitchell (1993) presented evidence using unemployment rates from 15 OECD countries that revealed high degrees of persistence following simulated 3 per cent negative output gap shocks. Mitchell (1993) concluded, "Convergence is slow relative to the actual frequency of shocks of this dimension experienced across the OECD block. Clearly, macroeconomic policy can be designed to minimize the costs of each shock (that is, reduce the output gaps) before the next shock impacts. A non-interventionist policy would see the impacts of previous shocks still 'substantially' in the system as the next shock arrives. Thus, the Okun losses would be magnified.

In recent years, there has been a major debate about the existence, measurement and time-varying nature of the so-called Non-Accelerating Inflation Rate of Unemployment (NAIRU). The NAIRU is closely tied to the Natural Rate Hypothesis (NRH), at least in terms of its policy brief. The NRH, a central pillar of orthodox, market-clearing theory, distinguishes between the long-term secular trend and the short-term (transitory) fluctuations in the economy. At best, aggregate demand management can only stabilise the short-term variations, but in the NRH it is usually considered to inhibit the *natural* tendencies of an economy (if shocked) to equilibrate, and ultimately only influences nominal magnitudes (that is, causes inflation). The evidence is mounting against the NRH and the usefulness of the NAIRU as a reliable indicator for policy (for example, Chang, 1997; Akerlof, Dickens, and Perry, 2000; Fair, 2000; Mitchell and Muysken, 2001; Mitchell, 2001). Further, evidence is now consolidating, which is not consistent with the view expressed by Blanchard (1981) and Kydland and Prescott (1980) that shocks to output are short-lived in their effects. A number of studies "have rejected the traditional view that output shocks have little or no permanent effect" (Campbell and Mankiw, 1987).[1]

A contrasting hypothesis is the hysteresis hypothesis (HH), which relates to path-dependence in dynamic systems (Cross, 1986; Mitchell, 1987; Franz, 1990). Franz (1990: 2) says "The long-run solution of such a system does not only depend on the long-run values of the exogenous variables (as usually) … [that is, under NRH models] … but also on the initial condition of each state variable." Buiter (1987: 24) expresses path-dependence as, "Where you get to is determined by how you get there." Accordingly, expansionary demand policy can permanently reduce unemployment at the cost of some inflation, the price level acceleration is finite as the economy adjusts to a new lower steady-state unemployment rate. The problem has been in distinguishing between the two hypotheses at an empirical level. Mitchell (1993) employed a unit root testing framework by stylising the NRH as inferring unemployment was within the trend-stationarity class of series and the HH as inferring that unemployment was difference-stationary. The problem of high levels of autoregressivity in the unemployment rates (and thus near-unit root status at least) makes it difficult to be conclusive given the lack of power of the tests.

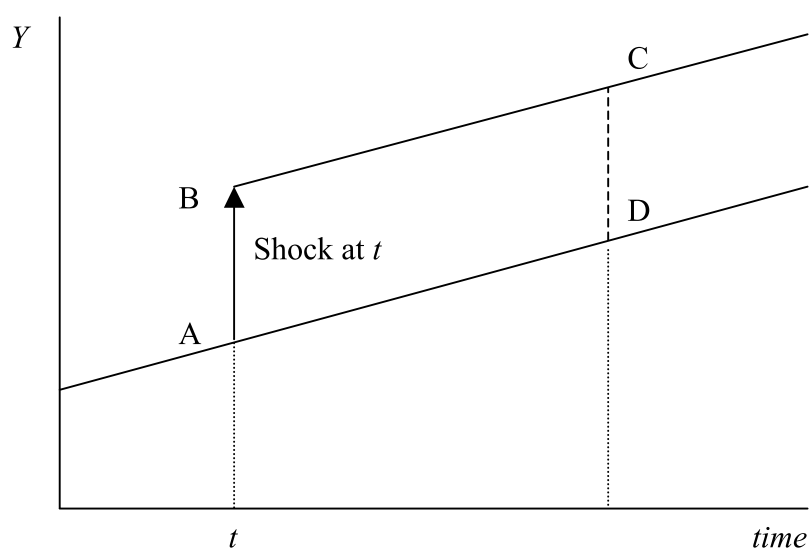
More important, for policy purposes, while the distinction between these hypotheses is clear in theory, on a practical basis the divide is somewhat blurred. Here the concept of unemployment persistence is important. In analytical terms, persistence is a special case of the NRH. An economy with strong persistence takes many periods to adjust back to equilibrium following a shock. So even if the NRH is a true model of the economy, persistence means that the effects of shocks have long memories and that short-term macroeconomic policy can be effective.

In this paper, a number of measures of persistence are employed to shed further light on the issue. The paper is laid out as follows. Section 2 briefly outlines the concept of persistence and distinguishes a unit root process from a persistent process. Section 3 presents a number of empirical measures of persistence. Section 4 briefly discusses the policy implications of the results. Concluding remarks follow.

2. The concept of persistence

The degree of persistence is the effect of a contemporaneous shock on the deviation of a time series process from its trend at some future date. The following diagram captures the types of processes classified according to their degree of persistence. We assume that the time series has some underlying trend. Following a unit shock at time *t* of AB magnitude, the time path of the adjustment is then shown according to the type of process. A process that resumes trend at D following the unit shock would be called Trend-stationary (or trend-reversion). The degree of persistence in this case will be governed by the combination of moving average and autoregressive components in the time series. A stationary time series with high degrees of autoregressivity will take many periods to settle back at D. A non-stationary process like a random walk (difference-stationary) would follow the BC path with the unit shock being permanent. Processes that have trajectories between CD and never return to the trend-reversion line are difference-stationary (DS) processes (containing at least one unit root) with long-run impacts being less than unity.[2] They are also exhibit stochastic non-stationarity.

Figure 1 Difference-stationary and Trend-stationary processes and persistence



The time path that these persistent DS processes take depends on the other roots in the process. An interesting case arises where the AR process is persistent but does not have a unit root. Take the case where the A(L) polynomial in an AR(2) process is $(1 - L + 0.25L^2)$, which means the AR(1) coefficient is unity. This process has two roots (both equal to 0.5) and is thus trend-stationary. The initial impact to a unit shock is 100 per cent but it decays to it previous pre-shock value within 15 periods. Simple near-unit root AR(1) processes with autoregressive coefficients of, say 0.98 will revert to their pre-shock values but are highly persistent (in this case the reversion takes 50 periods). The general conclusion is that if a process has a unit root then there will be some permanent effect on the level of the time series following a shock. However, persistence is common to unit root and near-unit root processes alike and can be measured independent of testing for the presence of a unit root. From a policy perspective it may be moot whether the process in fact formally tests for a unit root(s). Given the difficulty with the formal unit roots tests framework, measuring persistence directly is a useful exercise.

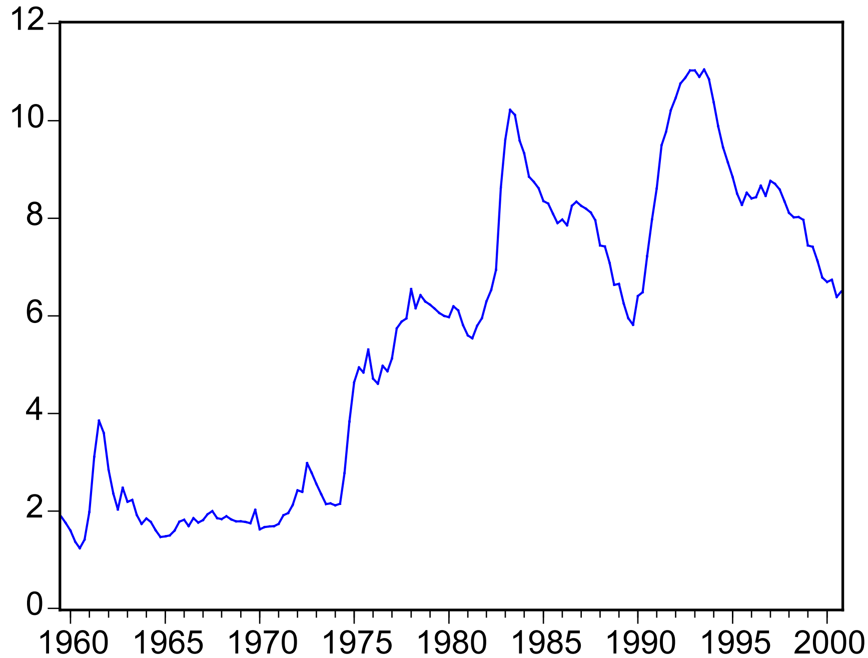
The interpretation of these persistence measures in terms of conventional depictions of macroeconomic dynamics is interesting. The overriding orthodox view is that there are two stylised facts about the business cycle: “First, fluctuations in output are assumed to be driven primarily by shocks to aggregate demand, such as monetary policy, fiscal policy, or animal spirits. Second, shocks to aggregate demand are assumed to have only a temporary effect on output; in the long run the economy returns to the natural rate. These two premises underlie many monetarist and neo-Keynesian theories” (Campbell and Mankiw, 1987: 876). If real variables are highly persistent it is clear that one or both of these facts is in error.

Real business cycle theory (see Nelson and Plosser, 1982) is based on a strong doubt about the first fact and instead claims that business shocks are mainly supply oriented. Recent evidence examining the empirical validity of real business cycle theory is not favourable (see Fair, 1994). Campbell and Mankiw (1987: 877) suggest that in relation to the second fact, models of nominal rigidities and/or misperceptions (for example, Lucas, 1973) would have to abandon the natural rate hypothesis in order to be reconciled with the evidence of high degrees of persistence. Two strands of thought emerge here. First, high levels of persistence are consistent with multiple equilibria models (Diamond, 1984) where aggregate demand shocks "move the economy between equilibria" (Campbell and Mankiw, 1987: 877). Secondly, aggregate demand shocks can persist if they instigate supply reactions (technological innovations).

3. Motivation

Figure 2 charts the Australian unemployment rate since 1959. The pattern is common to many OECD economies. Around the first oil shock in 1974, a major shock dislodged the Australian unemployment rate from its steady-state path of around 2 per cent. The shock saw the unemployment rate rise very quickly to a new level around 6 per cent. In the early 1980s, a further negative shock occurred, which pushed it above 12 percent. Over the strong growth period in the second half of the 1980s, the effects of this shock were steadily eroded. With the 1990s recession, the rate of unemployment rose sharply again, reaching heights not experienced since the Great Depression of the 1930s. Almost a decade later and after 37 quarters of positive GDP growth, the level has not yet recovered to that found at the previous peak.[3] The behaviour is consistent with a highly persistent process. In economics we are taught that to minimise costs is a sign of an efficient production process and a necessary condition for profit maximisation. How quickly the adherents to this paradigm forget when they move into the macroeconomic sphere. At this level, the dominant economic orthodoxy has since the mid-1970s cajoled policy makers to follow policies that have deliberately and persistently deflated their economies under the false pretext that the role of policy is to ensure the economy is operating at the natural rate of unemployment. The famous statement attributed to James Tobin that "it takes a lot of Harberger triangles to fill an Okun gap" has been ignored despite recent assessments that the costs of unemployment are huge (Mitchell and Watts, 1997; Watts and Mitchell, 2000; Harvey, 2000). The point is that when a highly persistent stock variable is shocked, the cost minimising strategy is likely to be one of intervention rather than leaving the market to sort out the dynamic adjustments. We examine this issue in a later section.

Figure 2 Australian unemployment rate, 1959:4 to 2000:4, per cent



A simple AR(1) regression for Australia for the period between 1959:4 and 2000:4 yielded the following results for the log of the aggregate unemployment rate (*t*-statistics are in parentheses):

$$\text{LUR} = 0.027 + \underset{(95.70)}{0.987}\ \text{LUR}(-1) \qquad R^2 = 0.98 \qquad \text{S.E. to Mean LUR} = 5.93\%$$

A negative shock to this series will persist for many quarters given the AR(1) coefficient of 0.987. For example, assuming a once-off, negative shock immediately pushes the unemployment rate up by 1 per cent and no further shocks are received. The half-life of this shock will be around 34 quarters.[4] The longer the time period before resolution is achieved, the larger are the cumulative costs to individuals and the economy in general.

What has been the experience of other OECD countries? Table 1 shows the AR(1) coefficients on the lagged unemployment rate for regressions, which also included a constant. The full samples periods are shown in column 2 and later columns should be interpreted within that constraint (see notes accompanying the Table). The results are indicative only as the AR(1) specification may not be the best representation of the underlying data generating processes. With that qualification in mind, the results reveal that, in general, the degree of persistence captured by the AR(1) coefficient has shifted over time. In most cases (where estimation was possible), there was a noticeable rise following the first oil shock in 1974. This was a sharp rise in Australia, Japan, Norway, Spain, the United Kingdom, and less so for Austria, Canada, Finland, and Germany. The rising trend was reversed somewhat in the 1990s for Australia, Finland, Japan, and Norway. In the cases of Italy and the United States the degree of persistence appeared to fall after the oil shocks and continued to do so over the 1990s. The results still convey high degrees of persistence in most countries (Italy, Japan, and Norway are probably exceptions).

Table 1 Shifting autoregressive parameters for OECD unemployment rates

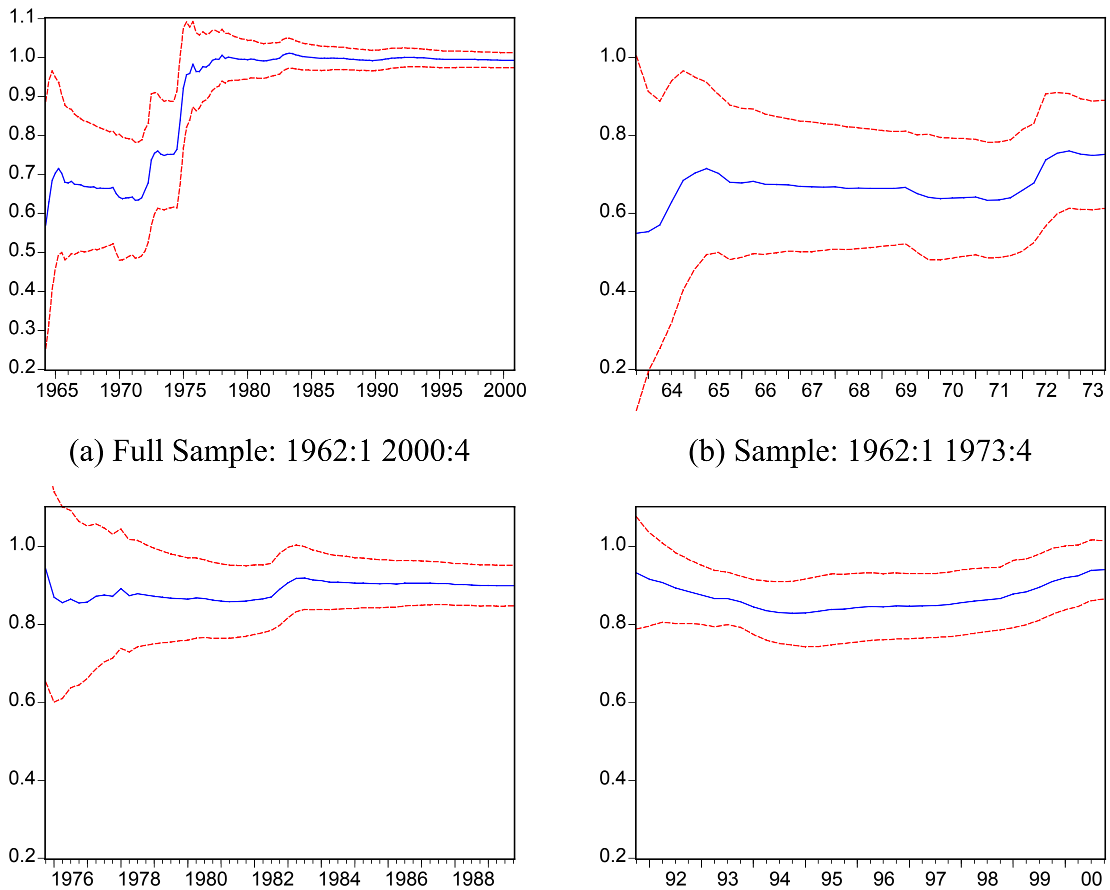
	Full Sample		Pre-oil	Post-oil	1970s	1980s	1990s
			61:4 73:1	74:1 89:4	70:1 79:4	80:1 89:4	90:1 00:4
Australia	1961:4 2000:4	0.970	0.760	0.900	0.980	0.963	0.940
Austria	1961:4 2000:4	0.995	0.974	0.987	0.825	0.942	0.911
Belgium	1980:4 2000:4	0.919				0.850	0.969
Canada	1961:4 2000:4	0.988	0.959	0.947	0.894	0.954	0.990
Denmark	1971:4 2000:4	0.961		0.824	0.956	0.845	1.026
Finland	1961:4 2000:4	0.984	0.918	0.961	0.990	1.098	0.917
France	1980:1 2000:4	0.947				0.927	0.973
Germany	1963:4 2000:4	0.989	0.909	0.927	0.970	0.918	0.967
Italy	1961:4 2000:4	0.986	0.937	0.982	0.918	0.954	0.808
Japan	1961:4 2000:4	1.005	0.697	0.885	0.949	0.922	1.008
Netherlands	1971:4 2000:4	0.948		0.888	0.895	0.939	1.037
Norway	1973:4 2000:4	0.954		0.929	0.554	0.959	0.972
Portugal	1985:1 2000:4	0.945				0.976	0.947
Spain	1966:1 2000:4	0.991	1.046	0.964	1.002	0.884	1.020
Sweden	1971:4 2000:4	0.991		0.975	0.910	0.992	0.922
Switzerland	1984:4 2000:4	0.989				1.039	0.936
United Kingdom	1961:4 2000:4	0.987	0.925	0.949	0.969	0.910	1.028
United States	1961:4 2000:4	0.988	0.949	0.939	0.855	0.984	1.018

Source: OECD Main Economic Indicators. The full samples are defined in column 2. In terms of the samples indicated in columns 4 to 8, starting dates for estimation are determined from the full sample starting dates. For example, for Switzerland, the 1980s starts at 1984:4. Missing values indicate no data for that sample. Some results were not reported because of too few observations (Switzerland and Portugal in the 1980s).

Focusing on Australia, we examine the nature of these shifts. In Figure 1 (and from data in the Appendix), we observe that the Australian economy experienced two major downturns in activity in the 1960s, both of which increased the unemployment rate significantly. In June 1960, the economy peaked and entered a 5-quarter downturn with a buildup in unemployment at the trough of 115 thousand over the decline. The recovery was very strong and within 3 quarters the unemployment rate had returned to its steady-state level of around 1.9 percent. Another severe shock in 1963 saw GDP growth fall from 2.1 per cent in the March quarter to –3.4 percent in June quarter. However, employment growth still met labour force growth in the one-quarter downturn and the economy quickly absorbed the build-up of unemployment.

The conclusion is that in the 1960s, negative shocks did not appear to persist. To examine this further, we estimated the AR(1) model using recursive least squares and plotted the time variation in the AR(1) parameter for four periods: the full sample, 1962:1 to 1973:4, 1974:1 to 1989:4, and 1990:1 to 2000:4. Figure 3 shows the stark results. Quadrant 3(a) shows that considerable instability in the coefficient occurred in the 1973-74 period. In 1974, the AR(1) coefficient dramatically rose towards unity and essentially remained at that new level for the rest of the sample. In other words, the major change in the degree of persistence occurred at that point. Prior to 1974, the degree of persistence was significantly lower. Quadrant 3(b) confirms that the sharp rises began after 1973:4. Quadrant 3(c) is interesting because the major disturbance to the coefficient appears to be cyclical (with the small rise and fall associated with the recession in the early 1980s). Overall, there was a slight upwards shift in the degree of persistence in this period largely driven by the 1982 recession. Quadrant 3(d) again confirms that recession appears to push the degree of persistence upwards. For the rest of the 1990s, the recovery in the Australian economy was associated with a steady drift upwards in the degree of persistence, although the level is somewhat lower than the level that followed the 1974 upheaval. Overall, the results are supportive of some hysteretic effects operating via the business cycle.

Figure 3 Australian unemployment rate recursive AR(1) coefficient



(a) Full Sample: 1962:1 2000:4

(b) Sample: 1962:1 1973:4

(c) Sample: 1974:1 1989:4

(d) Sample 1990:1 2000:4

Source: Recursive least squares regression of log unemployment rate on one lag and a constant. The plotted coefficient is the estimated AR(1) parameter for the samples noted.

However, it remains clear that the present behaviour of the unemployment rate in terms of reactions to shocks was strongly influenced by events in the short period during 1974. To explain the overall jump in persistence of unemployment one has to consider what happened in that period that was different. We tested for breaks in the periods shown around the troughs (see Appendix for details of timing). There was no statistically significant break associated with the 1982 and 1990 troughs.

However, the break in behaviour associated with the first oil shock is confirmed by the following regression, which added an interactive dummy variable (D75 = 1 after 1975:1 and 0 before) to the previous model (*t*-statistics are in parentheses):

$$LUR = 0.06 + \underset{(32.7)}{0.913}\, LUR(-1) + \underset{(3.05)}{0.11}\, D75$$

$R^2 = 0.99$ S.E. to Mean LUR = 5.30%

In seeking an explanation for the major break in the series during 1974, Mitchell (2001) argues that the 1974 recession was different to those that had occurred in the 1960s and also different to the two major downturns that followed it (1983 and 1991). The combination of deficient aggregate demand growth and rising labour productivity was the major reason for the rapid buildup in unemployment in that period.

4. Exploring sources of non-stationarity in OECD unemployment rates

Assessing the degree of persistence can also be approached more formally by examining the sources of non-stationarity in a time series. This was the approach taken by Mitchell (1993) who found an overwhelming degree of inertia in OECD unemployment rates. Two aspects of this analysis is updated in this section: (a) an examination of autocorrelation functions of the OECD unemployment rates in this study, and (b) a series of test of unit root hypotheses for the OECD unemployment rates. The data is from the quarterly OECD Main Economic Indicators. The log is used in every case.

4.1 Autocorrelation Functions

Table 2 shows the sample autocorrelations for each country in level form. Without any significant exception, the unemployment rates display a high degree of autoregressivity at lag one (the highest is 0.99, the lowest is 0.93), then slowly decay as the lag increases, with limited individual variations around this pattern.[5] These time series thus behave in a similar fashion to the ACF of a random walk (see Nelson and Plosser, 1982: 147).

Table 3 reports the ACFs for the first difference for each country. Most countries have a significant first lag and then their ACFs drop of rapidly at higher lags, a pattern consistent with stationarity. Some countries evade this trend with France, Switzerland and the United Kingdom being notable. To examine the possible sources of non-stationarity, linear filter $\left(\mu + \beta t\right)$ was put through each series and the ACFs computed for the 'de-trended' residuals of each series. The results are available on request but they are not consistent with stationarity. In this sense, the TSP alternative is not robust. The Dickey-Fuller ADF tests on these residuals were consistent with the evidence provided by the ACFs.

Table 2 Sample autocorrelations for *LUR*

		Lag					
Country	Period	1	2	3	4	5	6
Australia	1960:1 to 2000:4	0.98	0.96	0.92	0.90	0.88	0.87
Austria	1960:1 to 2000:4	0.99	0.97	0.96	0.94	0.92	0.90
Belgium	1979:1 to 2000:4	0.94	0.88	0.80	0.72	0.63	0.53
Canada	1960:1 to 2000:4	0.99	0.97	0.94	0.90	0.87	0.84
Denmark	1970:1 to 2000:4	0.96	0.91	0.85	0.79	0.74	0.69
Finland	1960:1 to 2000:4	0.98	0.96	0.93	0.89	0.85	0.81
France	1978:1 to 2000:4	0.95	0.89	0.84	0.79	0.74	0.69
Germany	1962:1 to 2000:4	0.98	0.96	0.93	0.91	0.88	0.85
Italy	1960:1 to 2000:4	0.99	0.97	0.96	0.94	0.92	0.90
Japan	1960:1 to 2000:4	0.98	0.96	0.93	0.90	0.88	0.85
Netherlands	1970:1 to 2000:4	0.95	0.89	0.82	0.76	0.69	0.62
Norway	1972:1 to 2000:4	0.96	0.93	0.89	0.84	0.80	0.75
Portugal	1983:2 to 2000:4	0.93	0.86	0.80	0.74	0.63	0.51
Spain	1964:2 to 2000:4	0.99	0.98	0.97	0.97	0.95	0.94
Sweden	1970:1 to 2000:4	0.98	0.95	0.92	0.87	0.82	0.78
Switzerland	1983:1 to 2000:4	0.98	0.96	0.92	0.88	0.83	0.77
United Kingdom	1960:1 to 2000:4	0.99	0.97	0.95	0.92	0.89	0.86
United States	1960:1 to 2000:4	0.97	0.92	0.86	0.79	0.72	0.65
Random Walk		0.95	0.90	0.85	0.81	0.76	0.70

Source: OECD Main Economic Indicators.
(b) From Nelson and Plosser (1982: 147), Table 2.

Table 3 Sample autocorrelations for ΔLUR

Country	Period	Lag 1	2	3	4	5	6
Australia	1960:1 to 2000:4	0.36	0.16	-0.13	-0.23	-0.13	-0.01
Austria	1960:1 to 2000:4	0.35	0.02	0.01	0.19	0.05	-0.04
Belgium	1979:1 to 2000:4	0.29	0.38	0.38	0.26	0.27	0.27
Canada	1960:1 to 2000:4	0.46	0.29	0.14	-0.02	0.00	-0.04
Denmark	1970:1 to 2000:4	0.49	0.10	0.07	0.11	-0.07	-0.19
Finland	1960:1 to 2000:4	0.39	0.27	0.41	0.14	-0.02	0.05
France	1978:1 to 2000:4	0.74	0.56	0.44	0.27	0.15	0.16
Germany	1962:1 to 2000:4	0.57	0.31	0.19	0.03	-0.01	-0.10
Italy	1960:1 to 2000:4	-0.10	0.14	0.14	-0.24	0.28	-0.15
Japan	1960:1 to 2000:4	0.07	0.05	0.21	-0.07	0.10	0.08
Netherlands	1970:1 to 2000:4	0.48	0.29	0.31	0.24	0.11	-0.01
Norway	1972:1 to 2000:4	-0.18	0.15	0.14	-0.21	0.08	-0.05
Portugal	1983:2 to 2000:4	-0.08	-0.08	0.06	0.49	0.05	-0.21
Spain	1964:2 to 2000:4	0.21	0.25	0.10	0.34	0.08	-0.01
Sweden	1970:1 to 2000:4	0.37	0.35	0.32	0.21	0.20	0.17
Switzerland	1983:1 to 2000:4	0.69	0.62	0.45	0.41	0.26	0.16
United Kingdom	1960:1 to 2000:4	0.66	0.45	0.29	0.15	0.09	-0.04
United States	1960:1 to 2000:4	0.65	0.37	0.20	-0.02	-0.09	-0.04
Random Walk		0.25	0.00	0.00	0.00	0.00	0.00

4.2 Formal Unit Root Tests

Various autoregressive representations can be used as the basis for unit root testing. For example, Perron (1988) defines three regression equations, which indicate an ordering of relevant hypotheses.

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

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

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

Equation (1.1) is stationary if $|\hat{\alpha}| < 1$, whereas if $\hat{\alpha} = 1$, the process has a unit root and is non-stationary (see Dickey and Fuller, 1979: 427, Equation 1.1). Equation (1.2) allows for fixed drift, μ^* (Dickey and Fuller, 1979: 428, Equation 2.1). Equation (1.3) provides the framework for testing: Hypothesis A a driftless random walk $\left(\widetilde{\mu}, \widetilde{\beta}, \widetilde{\alpha}\right) = \left(0, 0, 1\right)$, and Hypothesis B $\left(\widetilde{\mu}, \widetilde{\beta}, \widetilde{\alpha}\right) = \left(\widetilde{\mu}, 0, 1\right)$ (Dickey and Fuller, 1981: 1057, Equation 1.3) against a general alternative. We follow Perron (1988) and Mitchell (1993) in the sequence of hypothesis testing and start with Equation (1.3). This confronts the unit-root hypothesis against the obvious TSP alternative. To lessen any problems with serial correlation we use the Augmented Dickey-Fuller (ADF) regression model (Dickey and Fuller, 1981).[6]

The ADF regression format employed is

$$(1.4) \qquad y_t = \mu + \beta t + \alpha_1 y_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta y_{t-i} + e_t$$

To facilitate testing y_{t-1} is subtracted from both sides and regressed as:

$$(1.5) \qquad \Delta y_t = \mu^* + \beta^* t + \alpha_1^* y_{t-1} + \sum_{i=1}^{k} \gamma_i^* \Delta y_{t-i} + e_t$$

where $\alpha_1^* = (\alpha_1 - 1)$. The test becomes the straightforward test of $\alpha_1^* = 0$.

In terms of Hypothesis A, the Dickey and Fuller (1981) Φ_2-test is computed based on Equation (1.3). In terms of Hypothesis B, the Φ_3-test statistic (Dickey and Fuller, 1981) is computed based on Equation (1.3). We also calculate Φ_2^* and Φ_3^* based on the ADF regression. A range of ADF tests (for k = 0-4), the τ_τ-test from Fuller (1976).

4.3 Test results

The results are reported in Tables 4 and 5. Table 4 provides statistics for the unit-root null against the general alternative in Equation (1.3). Finland stands out as the only OECD economy with a significant ADF statistic. Table 5 presents the results for the joint tests. The Φ^*-statistics are based on the ADF regression (Equation 1.5), whereas the Φ-

statistics are from the Dickey-Fuller (DF) regression (Equation 1.3) The results vary due to the impact of the residual structure on the residual sum of squares in Equation (1.3).

Using the Φ_2-test of the joint hypothesis of a driftless random walk against the general alternative Equation (1.3) we can reject the null at the 5% level for Belgium, France, and the Netherland. However, Table 4 reveals significant fourth-order (except Japan, Italy, Norway and Portugal) serial correlation in the DF regression. In this case, the ADF regression is the preferable framework. Accordingly, the Φ_2^*-tests suggests that we can only reject the null for Finland and the Netherlands. The more restricted joint null of a random walk with fixed drift (Φ_3^*-test) once again adds Finland and the Netherlands as our two TS potentiates.

Table 4 Unit root regressions – *LUR* (model: $y_t = \hat{\mu} + \hat{\beta}t + \hat{\alpha}y_{t-1} + e_t$)

Country	T	$\hat{\mu}$	$t(\hat{\mu})$	$\hat{\beta}$	$t(\hat{\beta})$	$\hat{\alpha}$	$t(\hat{\alpha})$	$\chi^2(4)$
Australia	163	0.03	1.91	0.0002	0.54	0.97	-1.1951	31.94
Austria	163	0.01	0.93	0.0004	2.59	0.97	-2.3039	28.95
Belgium	87	0.11	2.47	-0.0003	-1.92	0.97	-1.3496	14.80
Canada	163	0.03	1.03	0.0000	0.28	0.98	-0.9178	40.01
Denmark	123	0.10	3.01	-0.0004	-1.14	0.98	-1.1932	31.93
Finland	163	0.02	1.25	0.0005	1.41	0.96	-1.6807	54.65
France	91	0.08	3.41	-0.0004	-2.78	0.99	-0.7490	43.42
Germany	155	0.02	0.71	0.0003	0.60	0.98	-1.1709	56.46
Italy	163	0.12	2.55	0.0005	2.34	0.92	-2.5175	17.95
Japan	163	0.00	-0.46	0.0006	3.24	0.94	-2.5963	10.34
Netherlands	123	0.11	5.41	-0.0008	-4.37	0.98	-1.5478	15.01
Norway	115	0.01	0.28	0.0008	1.37	0.92	-2.0847	14.05
Portugal	70	0.19	1.46	-0.0007	-1.26	0.94	-1.3833	22.65
Spain	146	0.04	2.91	-0.0005	-1.49	1.01	0.6967	22.52
Sweden	123	0.02	0.88	0.0001	0.28	0.98	-1.0035	31.49
Switzerland	71	0.13	1.10	-0.0009	-0.94	1.00	0.0590	40.14
United Kingdom	163	0.02	1.66	-0.0003	-1.92	1.01	0.6607	68.86
United States	163	0.03	0.90	-0.0001	-1.19	0.99	-0.6622	74.49

Based on these joint Φ tests, the majority of OECD countries examined appear to have DS unemployment rates. The $\tau(\hat{\alpha})$ test from Fuller supports this contention and emphasised the distinct behaviour of Finland.

Table 5 Dickey-Fuller joint hypothesis tests - *LUR*

Country	T	$\tau(\hat{\alpha})$ [b]	Φ_2 [c]	Φ_3 [d]	Φ_2^*	Φ_3^*
Australia	163	-1.33	1.31	1.26	0.98	1.06
Austria	163	-1.92	2.41	3.47	1.41	1.85
Belgium	87	-3.04	5.10	6.54	4.01	5.94
Canada	163	-2.29	0.38	0.56	1.75	2.63
Denmark	123	-1.80	3.11	3.87	1.63	2.34
Finland	163	-4.05	1.54	1.42	5.85	8.37
France	91	-1.36	16.32	19.01	2.62	3.72
Germany	155	-2.49	2.55	1.33	2.52	3.18
Italy	163	-2.24	2.32	3.17	2.02	2.56
Japan	163	-2.33	4.63	5.52	3.17	3.43
Netherlands	123	-2.68	11.06	15.72	4.98	7.36
Norway	115	-2.07	1.59	2.26	1.67	2.32
Portugal	70	-2.37	1.05	1.17	2.14	2.81
Spain	146	-0.97	3.56	2.12	2.34	2.57
Sweden	123	-3.13	0.82	0.71	3.30	4.93
Switzerland	71	-2.65	0.84	0.88	2.36	3.54
United Kingdom	163	-1.35	2.11	2.78	1.47	2.02
United States	163	-1.78	0.78	1.09	1.21	1.64

(a) Φ_2 and Φ_3 are based on $y_t = \mu + \beta t + \alpha y_{t-1} + e_t$, whereas Φ_2^* and Φ_3^* are based on $y_t = \mu + \beta t + \alpha y_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta y_{t-i} + e_t$, with k = 4. The values of T are based on the model without higher-order terms.

(b) Critical values for $\tau(\alpha)$ (see Fuller, 1976: 381, Table 8.5.2):

T =	80	100	120
5% level	-3.47	-3.45	-3.44

(c) Critical values for Φ_2 and Φ_2^* (Dickey and Fuller, 1981: 1063, Table v):

T =	80	100	120
5% level	5.03	4.88	4.86

(d) Critical values for Φ_3 and Φ_3^* (Dickey and Fuller, 1981: 1063, Table vi):

T =	80	100	120
5% level	6.59	6.49	6.47

We also examined the Phillips-Perron tests, which take into account the residual correlation in the ADF regressions and make appropriate adjustments. We concluded that there was no appreciable difference in the results using the Phillips-Perron framework compared with the ADF results reported. Full results are available from the author. A more diverse array of tests in available in Mitchell (1993), who considered the segmented trend models of Rappoport and Reichlin (1988) and Perron (1989). Exhaustive tested failed to establish any empirical credence for these hypotheses. It did not appear that the failure to reject the unit-root hypothesis was largely due to mis-specification of the original regression (that is, a disregard for any breaks in trend).

In conclusion, the formal unit root tests, while suffering from low power, point to high degrees of persistence in the unemployment rates examined. Blough (1988) argues that in small samples (especially frequently sampled data), trend-stationary processes are virtually observationally equivalent to DS processes with moving average errors (with roots close to minus one). Combining this knowledge with the evidence that at least the unemployment rates in the OECD countries examined are highly persistent, the results of the study provide further evidence for the mounting case that cyclical shocks can have long term effects on the unemployment rates in many OECD countries. We tentatively conclude that with the exception of Finland and perhaps the Netherlands the remaining OECD countries behave consistently with integrated processes of order one and are hence stochastic non-stationary over the sample period. In terms of our theoretical introduction, juxtaposing the NRH with the HH, this evidence is more consistent with the widespread presence of hysteresis across the OECD block than it is with the universality of the NRH. Of-course, the NRH hypothesis says nothing specific about the adjustment horizon following a deflationary period. So highly persistent processes could ultimately be consistent with the NRH.

5. Measuring Persistence

It thus becomes important to try to quantify the extent of persistence independent of the issue of whether the processes are TSP or DSP. In this section, we compute measures of univariate persistence that attempt to quantify the concept we outlined in Section 2. The estimates quantify the degree of persistence to a shock without attempting to identify the causes of the shocks. The degree of persistence indicated provides policy makers with a guide to the likely costs of non-intervention when a negative shock occurs. There are various approaches in the literature to measuring the degree of persistence. In this section we employ three such measures: (a) The ARMA(p,d q) measures proposed by Campbell and Mankiw (1987), (b) The Variance Ratio method proposed by Cochrane (1988), and (c) OLS and median-unbiased measures attributable to Andrews (1993).

5.1 ARMA measures

We begin by modelling the OECD quarterly unemployment rates in first difference form as ARMA(p,1,q) processes. The general model in logs is:

$$(1.6) \quad A(L)\left(\Delta y_t - \mu\right) = M(L)\varepsilon_t$$

where μ is the mean of the change in unemployment, $A(L) = 1 - \sum_{i=1}^{p} \alpha_i L^i$ and $M(L) = 1 + \sum_{i=1}^{q} m_i L^i$. The conditions for stability and invertibility (the respective roots of the polynomials are outside the unit circle) are tested before the persistence measures are computed.

Pesaran and Pesaran (1997: 225) argue that the "most satisfactory overall measure of persistence is the value of the spectral density of the first-differences of the series evaluated at zero frequency, and then appropriately scaled." For the ARMA(p,1,q) process the spectral density is given as:

$$(1.7) \quad f_{\Delta x}(0) = \frac{\sigma^2}{\pi}\left(\frac{1+\theta_1+\theta_2+\cdots+\theta_q}{1-\phi_1-\phi_2-\cdots-\phi_q}\right)^2$$

In this vein, Campbell and Mankiw (1987) propose a measure of persistence, which we denote *PCM*, which is derived by setting the lag operator equal to 1 and forming the ratio of the moving average and autoregressive polynomials. It is given as:

$$(1.8) \qquad PCM = \frac{M(1)}{A(1)} = \frac{1 + m_1 + m_2 + \ldots + m_q}{1 - \alpha_1 - \alpha_2 - \ldots - \alpha_p}$$

which is the restricted moving-average model of the process, with the lag operator L being set to unity. Computationally the Campbell and Mankiw (1987) PCM measure is simple but is sensitive to the order of the $A(L)$ and $M(L)$ polynomials identified. To reduce the sensitivity of the measure to the identification of the Box-Jenkins model, Pesaran and Samiei (1991) proposed an average measure of the PCM. In other words, we estimate a series of ARMA models (for $p = 1\ldots P$, and $q = 1 \ldots Q$) with the orders being sufficiently large to capture the dynamics of the data. We the calculate the PCM measures for each model as above and then take their arithmetic mean. For each country we estimated up to ARMA(6,1,6) and computed the PCM and averaged the 20 results.

To interpret the measure we note that a pure trend less stationary process would have a PCM measure of zero, while a pure random walk would yield a value of unity. The higher the value the higher is the degree of persistence. When the *PCM* is above unity, then we conclude that shock magnification occurs. Table 6 presents the *PCM* computations. Some countries appear to recover from shocks very quickly – Norway, Italy, and Japan have very low levels of persistence using the *PCM* measure. This is supported by the data in Table 1. Other countries like France are locked into high degrees of persistence. We compare these results with the Variance ratio measures explained in the next sub-section.

5.2 Variance ratio measures

Cochrane (1988) proposed a measure of persistence that is related to the Campbell and Mankiw (1987) approach but is estimated nonparametrically. Cochrane uses the autocovariances of the differenced process instead of the moving average coefficients that Campbell and Mankiw employ.

Cochrane (1988) derives his measure using the sample autocorrelations:

$$(1.9) \quad \hat{\rho}_j = \left(\frac{T}{T-j} \right) \sum_{t=j+1}^{T} \left(\Delta y_t \Delta y_{t-j} \right) \Big/ \sum_{t=1}^{T} \left(\Delta y_t \right)^2$$

To derive Cochrane's persistence measure, $\hat{V}^k$, the first k of these sample autocorrelations are used and weighted according to the following equation:

$$(1.10) \quad \hat{V}^k = 1 + 2 \sum_{j=0}^{k} \left(1 - \frac{j}{k+1} \right) \hat{\rho}_j$$

Cochrane (1988) says that this measure is a ratio of the variance of (k + 1)-period differences of the time series to the variance of the one-period differences. Priestley (1982: 463) has shown that in terms of frequency domain analysis, the variance ratio is "an estimate of the normalized spectral density at frequency zero using a Bartlett lag window. In frequency domain analysis, the value of the standardised spectrum at the zero frequency indicates the long-run properties of the time series. High values suggest that deviations from trend are persistent. For example, a pure unit root process has an infinite zero frequency.

There are many issues involved in using this estimator. Campbell and Mankiw (1987) consider several and conduct a sensitivity analysis for the size of k. They conclude (1987: 874) that "First, the window size k must be at least 30, and preferably 40 or 50, if one is to be able to discriminate between these two processes. Second, there is severe downward bias in $\hat{V}^k$; for the random walk, the mean of $\hat{V}^k$, is approximately (T – k)/T rather than unity. And finally, there is a great deal of sample variation in $\hat{V}^k$, so one must be cautious in making inferences based on this estimator."

The Campbell and Mankiw (1987) measure can be approximately related to the $\hat{V}^k$ measure via the following adjustment:

$$(1.11)\quad \hat{A}^k = \sqrt{\frac{\hat{V}^k}{1-R^2}}$$

where the R^2 is estimated from the ARMA model that is used to compute the Campbell and Mankiw measure. It can be replaced by the following:

$$(1.12)\quad \hat{A}^k = \sqrt{\frac{\hat{V}^k}{1-\hat{\rho}_1^2}}$$

which means we use the square of the estimated first-order autocorrelation. Campbell and Mankiw (1987: 322) argue that as this underestimates the R^2 in all cases other than AR(1) processes, it has to be adjusted to give an unbiased estimate of their measure. They propose a bias correction factor of $(T - \mathrm{k})/T$ (the mean of the random walk with drift in finite samples). This provides a reference against unity for the computed measure. We do not employ this correction here but note that it would reduce the estimates shown.

Table 6 present the *PCM* measure as outlined in the previous sub-section, the comparable Cochrane Adjusted measure (see Equation 1.12), and the unadjusted Cochrane measure (see Equation 1.10). In addition, we present standard errors for the $\hat{V}^k$ to provide some guide as to how useful these measures might be. It is interesting that Norway and Japan all appear to be low persistence countries. This is consistent across the three measures and the standard errors are relatively lower. In some cases, the Cochrane measures systematically reduce the degree of persistence compared to the *PCM* measure (for example, Australia goes from a relatively high degree of persistence to being below unity). In others, there are notable increases in the computed degree of persistence using the Variance ratio approach. In general, the non-parametric approach is less accurate than the ARMA route (see Campbell and Mankiw, 1987, 1989).

Figure 4 shows the variance ratio measures derived as in Equation (1.10) but estimated via the standardised spectral density at zero frequency using the Bartlett lag window for each country with 2 standard error bands shown. The results are replicated in Table 4. The noticeable feature is the wide standard errors for most countries, which makes it hard to place much confidence in the precision of the Variance ratio methods.

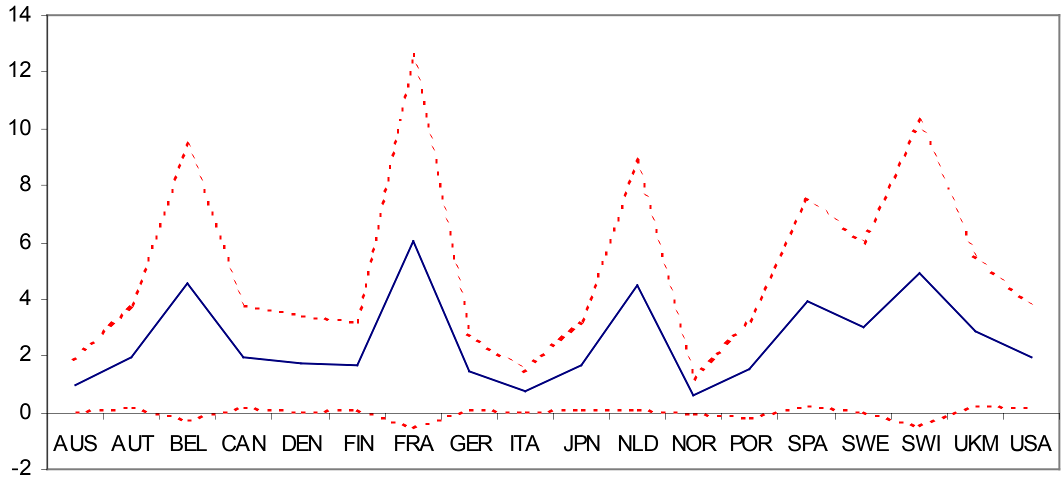
Table 6 Campbell and Mankiw (1987) and Cochrane (1988) measures of persistence

Country	Sample	Average PCM	Cochrane Adjusted	$\hat{V}^k$	Standard Errors for $\hat{V}^k$
Australia	1960:2 2000:4	1.57	1.06	0.97	0.449
Austria	1961:3 2000:4	1.56	1.50	1.98	0.911
Belgium	1979:2 2000:4	2.79	2.24	4.59	2.409
Canada	1960:2 2000:4	1.66	1.57	1.95	0.901
Denmark	1970:2 2000:4	2.01	1.52	1.75	0.854
Finland	1960:2 2000:4	2.12	1.41	1.69	0.779
France	1978:2 2000:4	4.48	3.66	6.04	3.268
Germany	1962:2 2000:4	2.11	1.49	1.48	0.672
Italy	1960:2 2000:4	0.89	0.86	0.74	0.342
Japan	1960:2 2000:4	0.43	1.30	1.68	0.772
Netherlands	1970:2 2000:4	1.97	2.41	4.49	2.194
Norway	1972:2 2000:4	0.79	0.78	0.59	0.300
Portugal	1983:3 2000:4	1.50	1.24	1.52	0.838
Spain	1964:3 2000:4	2.87	2.03	3.93	1.839
Sweden	1970:2 2000:4	2.96	1.87	3.03	1.478
Switzerland	1983:2 2000:4	3.73	3.06	4.91	2.693
United Kingdom	1960:2 2000:4	2.72	2.25	2.90	1.335
United States	1960:2 2000:4	2.23	1.82	1.93	0.890

$\hat{V}^k$ is estimated as the standardised spectral density at zero frequency using the Bartlett lag window with standard errors shown in the next column.

A reasonable conclusion is that these persistence measures while variable and open to contention, generally do not find low levels of persistence in the OECD unemployment rates examined.

Figure 4 Cochrane (1988) variance ratio V^k measure of persistence with standard errors



5.3 Median-unbiased measures

To further examine the issue of bias in the estimates of persistence, we examine the median-unbiased estimator (Andrews, 1993). Andrews developed the median-unbiased estimator to overcome bias in the standard least-squares based unit root testing framework. The standard Dickey-Fuller type tests are biased towards unity and have low power when faced with discriminating between unit root and near-unit root processes (see DeJong, Nankervis, Savin and Whiteman, 1992a, 1992b). In other words, even though we fail to reject the null of a unit root, we cannot be sure we have positive evidence in favour of the null. Andrews (1993: 141) argues that point and interval statistics have exact rather than asymptotic properties; exhibit smooth transition between trend-stationary processes and the difference-stationary processes (with or without drift); employ desirable initial conditions, and the median-unbiased estimator is unbiased in a fixed or variable-coefficient AR case. The limitation is that the median-unbiased estimator approach is only exact in the case of AR(1) processes (Rudebusch, 1992). Andrews approach allows us to construct an interval estimate of the AR coefficient, which helps us decide whether the failure to reject the unit root hypothesis is because the null is true or whether it is because we cannot accurately estimate the AR coefficient. For our purposes, the other useful aspect of the median-unbiased estimator approach is that it readily allows the computation of a persistence measure. Using the unbiased estimates of the AR coefficients we can compute cumulative impulse response (CIR) functions, which describe the "total cumulative effect of a unit shock on the entire future of the time series" (Andrews, 1993: 153) by number of periods. We can also compute the half life of a unit shock (HLS), which gives the number of periods taken until "the impulse response of a unit shock is half its original magnitude" (Andrews, 1993: 153). So we can compute the duration of shocks and the exact confidence interval pertaining to the estimated median duration of a unit shock.

The Andrews (1993) median-unbiased estimator is based on an AR(1) model such that:

$$(1.13) \quad Y_t = \tilde{\mu} + \tilde{\beta}t + \alpha Y_{t-1} + \varepsilon_t \qquad t = 1, \ldots, T$$

where $\tilde{\mu} = \mu(1-\alpha) + \alpha\beta, \quad \tilde{\beta} = \beta(1-\alpha), \quad \alpha \in (-1,1]$ and μ is a constant, *t* is a linear time trend, ε are the white-noise innovations to the process.

While the least squares estimator under classical assumptions is median-unbiased, the presence of the AR(1) parameter violates this property. The property is desirable because the median-unbiased estimator is impartial between underestimation and overestimation of the true value. This is particularly apposite where we are trying to distinguish between unit root and near-unit root processes. There is an equal chance of over- and underestimating the AR(1) parameter in the unit root regression. In other words, we have the same probability of choosing the true model as we have of choosing the wrong model.

However, Andrews (1993) shows how the least squares estimator can be corrected for median-bias. Assuming that $\hat{\alpha}$ is an estimate of the AR(1) parameter α whose median function m(α) is uniquely defined and with a true parameter $\forall \alpha \in (-1,1]$, then we can define the median-unbiased estimator of α as $\hat{\alpha}_U$:

$$(1.14)\quad \hat{\alpha}_U = \begin{cases} 1 & \text{if } \hat{\alpha} > m(1), \\ m^{-1}(\hat{\alpha}) & \text{if } m(-1) < \hat{\alpha} < m(1), \\ -1 & \text{if } \hat{\alpha} \le m(-1). \end{cases}$$

where m(-1) = $m(-1) = \lim_{\alpha \to -1} m(\alpha)$ and $m^{-1} : (m(-1), m(-1)] \to (-1,1]$ is the inverse function of m(.) that satisfies $m^{-1}(m(\alpha)) = \alpha$ for all $\alpha \in (-1,1]$. The interpretation is that as long as we have a function that generates the median value $\hat{\alpha}$ for each true value of α, then we can utilise the inverse function to derive the median-unbiased estimate of the true parameter. Andrews (1993) provides tables, which allow the median-unbiased estimator to be derived from the least squares estimator and the two-sided 90 per cent or one-sided 95 per cent confidence interval of α to be constructed. We term the resulting measure of persistence PA (after Andrews).

The impulse response function (IRF), an indicator of the degree of persistence of shocks to the time series, is given as (Andrews, 1993: 153):

$$(1.15)\quad IRF(\tau) = \alpha^{\tau} \qquad \tau = 0, 1, 2, \ldots$$

The *IRF* of a unit root process persists forever whereas in the case of a TS process, the *IRF* eventually dies, the duration of the death depending on the degree of persistence in the time series. This approach thus allows us to compare time series in terms of the respective persistence rather than be sidelined by the problems of discriminating between a unit root and a near-unit root TSP.

Andrews (1993) also computes two summary scalar measures of the *IRF*. The *CIR* and the *HLS* are given as:

$$(1.16)\quad CIR = \sum_{\tau=0}^{\infty} IRF(\tau) = \frac{1}{1-\alpha} \qquad HLS = \left|\log(1/2)/\log(\alpha)\right|$$

Median-unbiased estimators of each are derived by substituting the median-unbiased estimator of α for the least squares estimates. Exact confidence intervals are computed using the 0.05 and 0.95 quintiles for Equation (1.16).

The Andrews Procedure:

1. Run AR(1) regressions to obtain the least squares estimates of α.
2. Using the $T + 1$ sample size for Tables I to III (depending on the model chosen) in Andrews (1993), compute via extrapolation the median estimate, the 0.05 quantile and the 0.95 quantile. The last two computations provide the confidence intervals for a two-sided 90 per cent confidence interval.
3. Compute the median-unbiased estimators and exact confidence intervals for the IRF, CIR and HLS.

Table 7 contains the results of a comparison of the least-squares (LS) estimator of the AR(1) parameter to the median-unbiased estimator with 95 per cent confidence intervals shown as appropriate. The bias in the LS estimates are apparent and most countries display results consistent with a unit root in the unemployment rate. The estimates of the CIR and HLS point to very long shock persistence with standard errors in every case going to infinity. Denmark, France, the Netherlands, Norway and Portugal have HLS outcomes of around 3 years (2 for France). The French result is curious given the persistently high unemployment in that country. The other countries all have achieved relatively low unemployment rates over the 1990s.

Table 7 Least squares and median-unbiased estimates of persistence

Country	Estimator	α	IR(4)	IR(8)	IR(32)	CIR	HLS
Australia	LS	0.984	0.937	0.879	0.596	62.388	42.897
	MU	1.000	1.000	1.000	1.000	∞	∞
	CI	(0.977-1.000)	(0.909-1.000)	(0.827-1.000)	(0.467-1.000)	(42.550-∞)	(29.145-∞)
Austria	LS	0.995	0.982	0.964	0.865	221.037	152.864
	MU	1.000	1.000	1.000	1.000	∞	∞
	CI	(0.994-1.000)	(0.976-1.000)	(0.953-1.000)	(0.824-1.000)	(166.042-∞)	(114.744-∞)
Belgium	LS	0.946	0.801	0.642	0.170	18.538	12.500
	MU	0.992	0.970	0.941	0.784	132.210	91.294
	CI	(0.915-1.000)	(0.701-1.000)	(0.492-1.000)	(0.058-1.000)	(11.772-∞)	(7.808-∞)
Canada	LS	0.988	0.952	0.906	0.673	81.231	55.958
	MU	1.000	1.000	1.000	1.000	∞	∞
	CI	(0.980-1.000)	(0.923-1.000)	(0.852-1.000)	(0.527-1.000)	(50.474-∞)	(34.638-∞)
Denmark	LS	0.963	0.862	0.742	0.304	27.365	18.619
	MU	0.997	0.988	0.976	0.909	336.208	232.695
	CI	(0.945-1.000)	(0.799-1.000)	(0.639-1.000)	(0.167-1.000)	(18.363-∞)	(12.378-∞)
Finland	LS	0.991	0.963	0.928	0.741	107.256	73.997
	MU	1.000	1.000	1.000	1.000	∞	∞
	CI	(0.983-1.000)	(0.934-1.000)	(0.873-1.000)	(0.581-1.000)	(59.356-∞)	(40.795-∞)
France	LS	0.949	0.811	0.658	0.188	19.652	13.272
	MU	0.996	0.983	0.966	0.870	229.619	158.813
	CI	(0.918-1.000)	(0.710-1.000)	(0.504-1.000)	(0.065-1.000)	(12.197-∞)	(8.103-∞)
Germany	LS	0.988	0.953	0.908	0.678	82.937	57.140
	MU	1.000	1.000	1.000	1.000	∞	∞
	CI	(0.980-1.000)	(0.924-1.000)	(0.854-1.000)	(0.531-1.000)	(51.123-∞)	(35.088-∞)
Italy	LS	0.990	0.959	0.920	0.716	96.437	66.498
	MU	1.000	1.000	1.000	1.000	∞	∞
	CI	(0.982-1.000)	(0.930-1.000)	(0.866-1.000)	(0.561-1.000)	(55.911-∞)	(38.407-∞)

Notes: LS is the OLS estimator, MU is the median-unbiased estimator and CI are 95 per cent confidence intervals.

Table 7 (continued)

Country	Estimator	α	IR(4)	IR(8)	IR(32)	CIR	HLS
Japan	LS	1.007	1.000	1.000	1.000	∞	102.177
	MU	1.000	1.000	1.000	1.000	∞	∞
	CI	(1.000-1.000)	na	na	na	na	na
Netherlands	LS	0.962	0.858	0.736	0.294	26.634	18.112
	MU	0.996	0.984	0.968	0.879	249.177	172.370
	CI	(0.945-1.000)	(0.796-1.000)	(0.634-1.000)	(0.161-1.000)	(18.036-∞)	(12.152-∞)
Norway	LS	0.959	0.847	0.717	0.265	24.587	16.694
	MU	0.996	0.983	0.966	0.872	234.505	162.199
	CI	(0.941-1.000)	(0.785-1.000)	(0.616-1.000)	(0.144-1.000)	(17.040-∞)	(11.461-∞)
Portugal	LS	0.967	0.874	0.763	0.339	30.088	20.507
	MU	1.000	1.000	1.000	1.000	∞	∞
	CI	(0.947-1.000)	(0.805-1.000)	(0.649-1.000)	(0.177-1.000)	(18.982-∞)	(12.808-∞)
Spain	LS	0.993	0.972	0.944	0.794	139.339	96.236
	MU	1.000	1.000	1.000	1.000	∞	∞
	CI	(0.985-1.000)	(0.942-1.000)	(0.888-1.000)	(0.622-1.000)	(67.948-∞)	(46.751-∞)
Sweden	LS	0.984	0.938	0.879	0.597	62.534	42.998
	MU	1.000	1.000	1.000	1.000	∞	∞
	CI	(0.976-1.000)	(0.909-1.000)	(0.827-1.000)	(0.468-1.000)	(42.617-∞)	(29.192-∞)
Switzerland	LS	0.985	0.941	0.886	0.616	66.548	45.780
	MU	1.000	1.000	1.000	1.000	∞	∞
	CI	(0.977-1.000)	(0.9111-1.000)	(0.830-1.000)	(0.475-1.000)	(43.474-∞)	(29.786-∞)
UK	LS	0.989	0.958	0.918	0.710	94.055	64.847
	MU	1.000	1.000	1.000	1.000	∞	∞
	CI	(0.982-1.000)	(0.929-1.000)	(0.864-1.000)	(0.557-1.000)	(55.108-∞)	(37.850-∞)
USA	LS	0.986	0.947	0.897	0.647	74.071	50.995
	MU	1.000	1.000	1.000	1.000	∞	∞
	CI	(0.979-1.000)	(0.919-1.000)	(0.844-1.000)	(0.507-1.000)	(47.635-∞)	(32.670-∞)

Note: see Table 7 part 1 above.

6. Persistence and Buffer Stocks

The issue of persistence has relevance for the type of employment policy that might best be followed to achieve full employment. Cashin *et al* (2000) consider the persistence of shocks to world commodity prices. They find that for many of the series examined mean reversion of prices is an extremely long process, once the process is shocked. They consider a 60 month half-life (defined following Andrews, 1993) to be typical and suggest that it is "beyond which the cost (involving storage, financing, and output-reduction costs) of maintaining any stabilization scheme would like to be prohibitive."(Cashin *et al*, 2000: 201) They conclude that where "price shocks are highly persistent, then national or international arrangements to smooth price shocks will not be sustainable, and countries dependent on international trade in commodities affected by these long-lived shocks will need to adjust their macroeconomic and structural policies to conform with their new steady-state levels of national income, consumption, and wealth." (Cashin *et al*, 2000: 202)

What if we are talking about shocks to quantity variables, like unemployment? The unsustainability of price support schemes for agricultural commodities relates to the costs involved in supporting prices that are facing persistent price declines. In this paper, we have shown that negative shocks to the unemployment rates in the OECD countries examined persist for many periods. What are the implications of this in terms of the use of expansionary fiscal and monetary policy and public sector job creation to reduce the costs of this persistence (see Mitchell, 1987a, 1987b, 1993, 1994, 1996, 1998). Consider a buffer stock employment approach (see Mitchell, 1998). This is a reversal of the logic of price support schemes and avoids the problems identified by Cashin *et al (*2000).

In the case of persistent commodity price shocks, the issue is one of what constitutes a reasonable level of output when demand and prices are falling. The argument is not relevant when applied to available labour. We define full employment to be the state where there was no involuntary unemployment and that is ensured by a sufficient number of jobs to be available in relation to the supply of labour at the current money wage rates.

The reverse logic implies that if there is a price guarantee offered below the prevailing market price and a buffer stock of working hours constructed to absorb the excess supply at the current (private) market price, then we can generate full employment without encountering the problems of price tinkering. This approach has been termed the Job Guarantee (Mitchell, 2000)

Graham (1937) documents the ways in which the government might deal with surplus production in the economy. Graham (1937: 18) says, "The State may deal with actual or threatened surplus in one of four ways: (a) by preventing it; (b) by destroying it; (c) by 'dumping' it; or (d) by conserving it." Since the 1970s, when faced with an excess supply of labour, governments have adopted the "dumping" strategy via the NAIRU when the overall costs to the economy (and individuals) can be minimised using the conservation approach (Watts and Mitchell, 2000). Graham (1937: 34) notes,

> The first conclusion is that wherever surplus has been conserved primarily for future *use* the plan has been sensible and successful, unless marred by glaring errors of administration. The second conclusion is that when the surplus has been acquired and held primarily for future *sale* the plan has been vulnerable to adverse developments …

The distinction is important to the Job Guarantee approach. Commodity price support schemes are typically examples of storage for future sale and are not motivated to help the consumer of wool but the producer. The Job Guarantee policy is an example of storage for use where the "reserve is established to meet a future need which experience has taught us is likely to develop" (Graham, 1937: 35). Graham also analysed and proposed a solution to the problem of interfering with the relative price structure when the government built up the surplus. In the context of the Job Guarantee, this means setting a guaranteed wage *below* the private market wage structure, unless strategic policy in addition to the meagre elimination of the surplus was being pursued. For example, the government may wish to combine the JG policy with an industry policy designed to raise productivity. In that sense, it may buy surplus labour at a wage above the current private market minimum. In the first instance, the basic JG model with a wage floor below the private wage structure shows how full employment and price stability can be attained. While this is an eminently better outcome in terms resource use and social equity, it is just the beginning of the matter.

Graham (1937: 42) considered that the surplus should "not be pressed for sale until an effective demand develops for it." In the context of the Job Guarantee policy, this translates into the provision of a government job for all labour, which is surplus to private demand until such time as private demand increases.

Conclusion

In this paper we have examined the concept of persistence in OECD unemployment rates. The techniques employed have been selective but the general conclusions are similar. There is considerable persistence to shocks demonstrated. At the outset, we indicated we were motivated by a concern that non-intervention following a negative shock was a costly strategy when the process receiving the shock was highly persistent. To that end, we advocate an endogenous employment policy, which breaks the nexus between unemployment persistence and negative shocks. Mitchell (1998) has shown that this is the rational strategy for a government that issues fiat currency and wants to minimise the costs of flux and uncertainty in the economies subject to large demand shocks.

Appendix

Table A1 Dating Australian Business Cycles.

Business Cycle Reference Dates		Duration in Quarters			
		Contraction	Expansion	Cycle	
Trough	Peak	Trough from Previous Peak	Trough to Peak	Trough from Previous Trough	Peak from Previous Peak
na	June 1960				
September 1961	March 1963	3	5	na	9
June 1963	December 1965	1	10	7	11
June 1966	December 1973	2	30	12	32
June 1974	March 1982	2	31	32	33
June 1983	December 1990	5	30	36	35
June 1991	September 2000	2	37	30	39

Source: ABS AUSSTATS National Accounts. The methodology is described in the text. na indicates that the data was not available to compute the starting point of the cycle that concluded in June 1960

Table A2 Correspondence between GDP and Employment Peaks and Troughs

GDP Trough	GDP Peak	Employment Trough	Employment Peak
na	June 1960	na	December 1960 (q)
September 1961	March 1963	March 1961 (q)	*
June 1963	December 1965	*	*
June 1966	December 1973	*	June 1974 (q)
June 1974	March 1982	March 1975 (q)	January 1982
June 1983	December 1990	April 1983	July 1990
June 1991	September 2000	July 1991	August 2000

Source: Table A1. The employment peak/troughs from January 1982 were based on monthly labour force data (ABS, The Labour Force, Australia, 6203.0) while those preceding this data were based on the quarterly version of this survey (denoted (q)). * refers to the no intervening peaks or troughs of merit.

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[1] These studies include Campbell and Mankiw (1987, 1989), Cochrane (1988).

[2] Take a process with a A(L) polynomial ($1-0.5L-0.5L^2$). This still has a unit root but the second non-zero (negative root) reduces the degree of persistence. In this case, the level of y would in the long-run be 0.67 units higher following a unit shock but will oscillate towards this new equilibrium.

[3] In the Appendix we provide a brief historical account of business cycle behaviour in Australia.

[4] The half life is the time it takes for 50 percent of the shock to die. The time path of the variable following the shock is given by $y_t = y_0 + \alpha^t$ where y_0 is the steady-state value and α is the AR(1) parameter (in this case, 0.98) and t is a time period following the shock. When y_t returns to y_0 the shock has fully died.

[5] Compared to Mitchell (1993) the ACF values have risen slightly more or less uniformally.

[6] Phillips (1987) addressed the serial correlation issue by developing a non-parametric approach to eliminate the dependence of the asymptotic distribution of his modified test statistics on the correlation structure of the residuals.