DO THE PHASES OF THE BUSINESS CYCLE DIE OF OLD AGE?

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The paper re-examines the issue of duration dependence in the Australian classical and growth business cycles in light of the somewhat surprising results obtained recently by Cashin and Ouliaris (2004). In so doing the authors use the multinomial logit regime switching modelling approach of Layton and Smith (2003). The paper also represents an extension of the earlier work on the issue undertaken by Bodman (1998); the key extensions being that the issue is framed within an explicit established business cycle chronology, a leading index is also included within the analysis, and the growth cycle, in addition to the classical cycle, is considered. Strong evidence of duration dependence is found for periods of recession within the classical cycle and for both phases of the growth cycle. Moderate evidence of duration dependency is also found for periods of classical cycle expansion. However, the evidence in this regard is significantly reduced once movements in the leading index are included in the analysis with its movements exhibiting strong power in predicting the termination of classical business cycle expansions. For growth cycles, duration dependence symmetry is found across both phases of the cycle.

I. Introduction

The current period of expansion being experienced in the classical Australian business cycle is now the longest of the post-war era. Up until the time of analysis (early 2004) it had achieved a duration of 149 months and with favourable economic conditions forecast for some time to come it will now significantly exceed the record length of the 1960s (into the early 1970s) economic expansion of 153 months. A question that naturally arises from such a lengthy period of expansion is whether the expansion is more likely to terminate as a result of being so long; i.e. whether the phases of the business cycle are characterised by positive duration dependence?

Arguments in favour of such duration dependency are often espoused in the popular press and by market analysts. Such commentary suggests that the duration of business cycle phases tends to cluster around some average length such that, the older the phase, the greater the probability that it will end. Burns (1969) argued that restrictive forces inherent within the economic system gradually, but insistently, cause the phases of the business cycle to terminate over time. Haberler (1937) draws an analogy between the rationale for duration dependency in phases of the business cycle and the presence of duration dependence in human mortality whereby assorted stresses and strains accumulate over time thereby increasing the probability of death. Additional support for such a characterisation is found within the multiplier-accelerator and inventory system business cycle models of Samuelson (1939) and Metzler (1947), and the structural propagation mechanism of Frisch (1933).

Fisher (1925) however argues that business cycles represent random walks such that any historical information concerning the cycle has no predictive value. This conceptualisation views business cycles as random strings of increases and decreases similar to the phantom luck

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perceived by gamblers at a casino. Such a view assumes that historical information is irrelevant for determining the phase of the business cycle and therefore suggests that duration does not influence the probability of a phase switch. The phases are therefore considered to be duration independent and the probability of a phase switch is assumed to be constant over time.

Recent empirical investigations of duration dependence within the Australian business cycle have produced mixed results. Bodman (1998) employs a variant of the Markov regime switching model of Hamilton (1989) and finds significant evidence of duration dependence during periods of recession but not for periods of expansion. Cashin and Ouliaris (2004) employ non-parametric methods and somewhat surprisingly fail to find any significant evidence of duration dependence across either phase of the Australian business cycle. The results of Cashin and Ouliaris (2004) appear to be at odds, not only with Bodman (1998), but also with studies of the US business cycle. US studies provide quite strong evidence supporting duration dependence in recessions (Sichel, 1991; Diebold and Rudebusch, 1991; Durland and McCurdy, 1994; Layton and Smith, 2003; and Zuehlke, 2003). The more recent studies of Layton and Smith (2003) and Zuehlke (2003) also provide some – albeit weaker – evidence supporting duration dependence in US expansions.

A related area of research is whether the phases of the growth cycle exhibit duration dependence. Growth cycles rose to prominence from the late 1960s due to the relative mildness of post-WW2 macroeconomic fluctuations which led a number of economic commentators at the time to question the continuing existence of a classical business cycle (e.g. Bronfenbrenner, 1969). Fluctuations within the rate of economic growth have continued to occur and, as such, growth cycles capture these through altering the definition of the business cycle by defining it in terms of a relevant measure of aggregate economic activity adjusted for its long run trend rate of growth.

Intuitively, it would appear that the phases of the growth cycle should exhibit duration dependence by virtue of the de-trending procedures applied when defining the cycle. The statistical process of fitting a representative trend to a set of data, removing it, and then considering the resulting residual fluctuations would logically suggest there would be in evidence a strong tendency for the periods of time above and below trend to terminate as the duration of the time in each growth phase lengthened. Such an argument is supported by Abderrezak (1998) who employed a parametric hazard function model and provided significant evidence – for a number of countries including Australia – of positive duration dependence across both phases of the growth cycle. More recently, however, again Cashin and Ouliaris (2004) used non-parametric methods and concluded that fast growth phases of the Australian growth cycle did not exhibit any evidence of duration dependence, whilst, quite surprisingly, finding that slow growth phases of the growth cycle exhibited evidence of *negative* duration dependence. Negative duration dependence in the growth cycle implies that the probability of staying in a below trend growth phase *increases* with the age of the phase, in contrast to declining as would seem more intuitive given the statistical manner (just described) in which a growth cycle is constructed. The Cashin and Ouliaris paper is the only empirical study – as far as the authors are aware – which suggests the possibility of negative duration dependence within the phases of the growth cycle. As such, further research would appear to be required to examine the issues.

This paper then has two main aims. Firstly, using the methodology of Layton and Smith (2003), to extend the previous work undertaken by Bodman (1998) by investigating the presence of duration dependency within an established chronology of the Australian classical business cycle and to consider the effects of changes in a leading economic index within the analysis. Secondly, in light of the results obtained by Cashin and Ouliaris, to investigate duration dependency within the Australian growth cycle.

The paper is organised as follows: Section II briefly discusses the differing empirical frameworks used in detecting the presence of duration dependence. Section III applies the framework of Layton and Smith (2003) to business cycle data for the classical and growth cycles of the Australian economy, with Section IV presenting conclusions.

II. Empirical Framework

Cashin and Ouliaris (2004) employ the non-parametric techniques of Brain and Shapiro (1983) to investigate duration dependence. This methodology seeks to test the durations of the completed phases of the business cycle for deviations from the null that the data are derived from an exponential probability distribution. The exponential distribution is hypothesised because it is characterised by a constant hazard function and hence suggests that the data exhibit duration independence.[1] Non-parametric techniques assume the ex-post observability of business cycle phases such that the duration data used within the tests can be defined. Cashin and Ouliaris (2004) define the Australian business cycle through using the Bry and Boschan (1971) algorithm applied to GDP.

Sichel (1991) outlined a number of criticisms associated with non-parametric techniques in the context of the current exercise. The most notable being the lack of power they exhibit in detecting duration dependence. Ohn, Taylor, and Pagan (2002) also argue that the null hypothesis of conformity to the exponential distribution, as tested by Cashin and Ouliaris (2004), is flawed because the duration data are discrete in nature and, as such, it is the geometric distribution which is more appropriate. Additionally, a number of authors have criticised the methodology of using a single macroeconomic indicator as a proxy for the business cycle (Boehm, 1998; Layton and Banerji, 2003). Adopting a single measure of the business cycle fails to capture the many activities that constitute the complex phenomena that is the business cycle.

Bodman (1998) investigates duration dependence using the regime switching framework of Hamilton (1989) which seeks to model a time series experiencing a number of different phases. The business cycle, which Bodman measures through a single macroeconomic proxy (either GDP or the unemployment rate), is often conceptualised as experiencing two distinct phases, i.e. expansions and contractions, which are summarised by the discrete random variable, S_t. Hamilton (1989) suggested that a comprehensive model for such a series should not only include a probability rule that characterises the differing phases but also a description of the probability rule that governs the changes between the distinct phases.

A simple probability rule to characterise each of the regimes is the normal density function. The normal density function governs the likelihood of observing various observations of the time series from each phase by adopting differing means and variances across each phase. Hamilton (1989) proposes a first order Markov chain to model the transition between the distinct phases. Such a model suggests that the probability of being in a particular phase of the business cycle depends on the past only through the most recently observed phase:

$$P(S_t = j \mid S_{t-1} = i, S_{t-2} = k, \ldots) = P(S_t = j \mid S_{t-1} = i) = p_{ij}$$

where $S_t = \begin{cases} 1, & \text{expansion} \\ 2, & \text{recession.} \end{cases}$

[1] Refer to Kiefer (1988) for a more detailed account of non-parametric techniques.

The transition matrix describing the evolution of S_t is thus given by

$$\mathbf{P} = \begin{bmatrix} p_{11} & 1 - p_{22} \\ 1 - p_{11} & p_{22} \end{bmatrix}$$

Within the transitional matrix $\mathbf{P}$, p_{11} denotes the probability of remaining in an expansion from period $t-1$ to period t, and p_{22} is the probability of remaining in a recession from period $t-1$ to period t. Given that the transition probabilities for each of the phases after a given phase has been observed sum to unity, the off diagonal elements are simply: $p_{12} = 1 - p_{11}$ denoting the probability of switching from an expansion to a recession; and $p_{21} = 1 - p_{22}$ denotes the probability of switching from a recession to an expansion. The original model proposed by Hamilton (1989) assumes that the transition probabilities are constant, and are therefore simply included within the set of parameters to be estimated.

The regime switching model provides a framework for testing for duration dependence through relaxing the restrictive assumption regarding constant transition probabilities. The transitional probabilities can alternatively be modelled as time varying functions of whatever determinants, X_{t-1}, may be considered potentially relevant to driving them:

$$P_{iit} = \frac{1}{(1 + \exp(-\gamma_i' X_{t-1}))}; \quad i = 1,2;$$

where $X_{t-1} = (1 \cdot x_{1,t-1}, x_{2,t-1}, \ldots, x_{k-1,t-1})'$; $\gamma_i = (\gamma_{i0}, \gamma_{i1}, \ldots, \gamma_{i,k-1})'$ and $k-1$ is the number of determinants of the transition probabilities.[2] Filardo (1994) and, in the case of Australia, Layton (1997), suggested that composite leading indicator indexes represent possible determinants of the transition probabilities. Durland and McCurdy (1994) allow the transition probabilities to vary over time according to the number of periods the system has been in a particular state, i.e. phase duration. Therefore the extent of duration dependence can be analysed through testing the significance of phase duration, d_{t-1}, as an explanatory variable of the transition probabilities.

The difficulty with including duration as an explanatory variable of the time varying transitions is that the model regards phase duration as unobservable. One of the key features of the model is that it assumes the specific phases of the time series being modelled are unobservable and, as such, infers the probability of each observation coming from a particular phase. Phase duration becomes dependent upon the model's inferences regarding the likelihood of the series being in a particular phase at each point in time. The values of the duration explanatory variable become part of the estimation process itself and therefore the estimation of the model requires a consideration of all of the possible phase patterns over the sample period being studied.

Consideration of all of the possible phase patterns gives rise to an exponentially expanding range of possibilities, the number of these patterns is 2^T, where T is the sample size and, clearly, estimation of the model becomes infeasible. To reduce the magnitude of this problem Durland and McCurdy (1994) suggest arbitrarily truncating the duration variable at some maximal value D, above which the transition probability is assumed to remain constant. Whilst this facilitates estimation, the choice of the value of D becomes crucial and could potentially very seriously affect the reliability of any resulting implied conclusions as to the statistical significance or otherwise of duration as an explanation of phase changes.

[2] The functional form, $1/(1 + \exp(x))$ represents the logistic function which is one of several that could be used which ensure the estimated transition probabilities at each time period are bounded between zero and one.

The multinomial regime switching logit model of Layton and Smith (2003) represents a regime switching framework for directly modelling the transition probabilities assuming the ex-post observability of business cycle phases. The approach adopted by Bodman (1998) would appear to be most useful in situations where knowledge of the phases for the series being modelled is unavailable. For business cycles, however, there appears to be a number of useful business cycle chronologies that can be used to define its phases. For the US, Layton and Smith (2003) incorporated this information within the modelling process and therefore eliminated the uncertainty associated with the occurrence of phase switches and hence the value of the duration variable at each point in time.

Incorporating such information within a regime switching model has the potential to significantly reduce the complexity of the model and increase the precision of the estimates of the various parameters of the model. Additionally, including such information allows for the issue of duration dependence to be framed in terms of whatever duration dependence is evident within the particular business cycle chronology regarded as being appropriate.

The multinomial regime switching logit model of Layton and Smith (2003) incorporates a first order Markov chain to directly model the transition between observed phases of the business cycle, as defined by some established explicit business cycle chronology. The assumed ex-post observability of the phases of the business cycle means the dependent variable of the model representing the business cycle phase changes is observable with the model's parameters being obtained by maximising the likelihood of the occurrence of the observed phase changes.

In this set-up the dependent variable represents the possible phase changes which could occur between any two consecutive periods. Therefore, even though there are only two possible alternative phases of the business cycle that may be observed at each point in time, there are four possible outcomes that can occur over any two consecutive periods:

1. The business cycle stays in an expansion: $S_{t-1} = 1$ and $S_t = 1$.
2. The business cycle experiences a turning point whereby it switches from a period of expansion into a period of recession: $S_{t-1} = 1$ and $S_t = 2$.
3. The business cycle experiences a turning point whereby it switches from a recession into a period of expansion: $S_{t-1} = 2$ and $S_t = 1$.
4. The business cycle stays in a recession: $S_{t-1} = 2$ and $S_t = 2$.

Therefore the dependent variable within the model can be considered to be multinomial in nature with four possible values. Four dummy variables, each with N observations, are used to represent which of the four possible outcomes outlined above actually occurred at each point in time and hence define the dependent variable of the regime switching multinomial model, viz:

$$h_t^A = \begin{cases} 1 & \text{if } S_t = 1 \text{ and } S_{t-1} = 1 \\ 0 & \text{otherwise} \end{cases}$$

$$h_t^B = \begin{cases} 1 & \text{if } S_t = 2 \text{ and } S_{t-1} = 1 \\ 0 & \text{otherwise} \end{cases}$$

$$h_t^C = \begin{cases} 1 & \text{if } S_t = 1 \text{ and } S_{t-1} = 2 \\ 0 & \text{otherwise} \end{cases}$$

$$h_t^D = \begin{cases} 1 & \text{if } S_t = 2 \text{ and } S_{t-1} = 2 \\ 0 & \text{otherwise} \end{cases}$$

At each point in time, t, only one of the dummy variables, h_t, can take the value of one, indicating which specific outcome occurred at that particular time period.

Thus, using the logistic function specification, we have the probability of staying in phase i ($i = 1,2$) given as

$$P(S_t = i \mid S_{t-1} = i, X_{t-1}) = (1 + \exp\{-(\alpha_i + \delta_i d_{t-1} + \beta_i' Z_{t-1})\})^{-1}$$

where Z_{t-1} is a column vector of selected leading economic indicators (with β_i representing the vectors of associated parameters), d_{t-1} is the duration of the current expansion or recession up to period $t-1$ (with associated parameters, δ_i) and defined as

$$d_{t-1} = \begin{cases} d_{t-2} + 1 & \text{if } S_{t-1} = S_{t-2} \\ 1 & \text{if } S_{t-1} \neq S_{t-2} \end{cases},$$

and the use of the logistic transformation, $(1 + \exp\{-x\})^{-1}$, guarantees the estimated probability will take a value between zero and one.

Parameter estimates of the model are obtained by maximum likelihood estimation. The log likelihood function is defined as:

$$\begin{aligned} LL(h;\theta) = \sum_{t=1}^{T} & h_t^A \log(\Phi(\alpha_1 + \delta_1 d_{t-1} + \beta_1' Z_{t-1})) + h_t^B \log(1 - \Phi(\alpha_1 + \delta_1 d_{t-1} + \beta_1' Z_{t-1})) \\ & + h_t^C \log(1 - \Phi(\alpha_2 + \delta_2 d_{t-1} + \beta_2' Z_{t-1})) + h_t^D \log(\Phi(\alpha_2 + \delta_2 d_{t-1} + \beta_2' Z_{t-1})). \end{aligned}$$

where Φ simply denotes the logistic functional form.

Through maximising the value of the log likelihood, estimates of the model's parameters can be obtained. Given the logistical functional form, if the coefficient on the duration variable is negative (positive), then the phase exhibits positive (negative) duration dependence. Alternatively, if the coefficient on the duration variable is insignificantly different from zero then the phase would be considered to exhibit duration independence. The standard likelihood ratio test gives an overall sense of the presence of duration dependency within the phases of the business cycle by comparing the estimated likelihood from the unrestricted model allowing for phase duration dependency with the restricted model incorporating duration independence.

III. Empirical Results

Data used in this study are monthly and are obtained from the business cycle chronology of the Economic Cycle Research Institute (ECRI), spanning the available post-WW2 period through until December 2002.[3] These data specify the phases of the business cycle and therefore define the value of the phase state variable, S_t, and thus the value of the duration variable, d_t, at all points in time:

$$d_{t-1} = \begin{cases} d_{t-2} + 1 & \text{if } S_{t-1} = S_{t-2} \\ 1 & \text{if } S_{t-1} \neq S_{t-2} \end{cases}$$

This chronology represents one of a number of chronologies that could be adopted. Other methodologies that could have been adopted to identify the business cycle chronology include sequencing rules, such as that which identifies a peak (trough) as the quarter before two consecutive quarters of negative (positive) GDP growth, or perhaps the chronology resulting from the application of, say, the turning point algorithm of Bry and Boshcan (1971) to some series regarded as suitably representing aggregate economic activity. The advantage in adopting the established chronology of ECRI is that the methodology used by this organisation follows the original

[3] The turning points for both the classical and growth cycles are provided in Table I.

NBER procedures for identifying phases of the business cycle and, as such, makes reference to a range of economic time series in measuring the cycle, as well as adhering to the various censoring rules as determined by the NBER, e.g. no individual phase is shorter than five months in duration, and no whole cycle is less than 15 months in duration. This chronology generally enjoys a wide acceptance as representing useful information in identifying the business cycle.[4]

Before outlining the empirical results and conclusions, the arguments made by Pagan (2005) in a very recent and important CAMA Working Paper need to be noted.[5] Pagan points out that a constructed phase state variable such as S_t derived from a country's business cycle chronology using either NBER-type methods or the use of some other turning point algorithms are very likely to be serially correlated. This is readily appreciated by considering that business cycle phases extend for quite some months (especially expansions but also, to a lesser extent, recessions), meaning that S_t may equal one (or two) for quite some number of consecutive periods.

This is an important observation which has implications for estimation and inference for all empirical investigations of a country's business cycle – including this paper – which use such constructed binary variables in one way or another, as either regressands or regressors.[6] In particular, as is well-known, *t*-ratios using conventional estimated standard errors can be exaggerated in the presence of serial correlation. Therefore, as far as Tables II–IV are concerned, whilst the robust standard errors are robust in respect of deviations from normality of the error term, the reader needs to note that there is the possibility that there may be some degree of underestimation of the standard errors of the models' estimated coefficients.

a) The classical cycle

The estimates for the model for the classical business cycle chronology assuming constant transition probabilities (Model 1) and the unrestricted model that allows the transition probabilities to vary over time according to current duration (Model 2) are produced in Table II.

Allowing the transition probabilities to vary over time according to phase duration significantly increases the explanatory power of the model for the Australian classical cycle. The value of the log likelihood increases from −45.01 to −41.61, representing a likelihood ratio statistic of 6.8 which is sufficiently large in comparison to the critical value at the five per cent level of significance of 5.991. This increase in the explanatory power of the model suggests that allowing the transition probabilities to vary according to phase duration provides a superior description of the data in comparison to the constant transition probability model and, as such, provides evidence in support of duration dependence.

The coefficients of the individual duration variables within the model also provide evidence of duration dependency. The duration variable across both phases of the business cycle displays

[4] The adoption of ECRI's chronology was also adopted in this study over other established chronologies for Australia as originally a range of other countries were also analysed in the manner described in the paper. Because ECRI produces chronologies for a large number of different economies including Australia, adoption of ECRI's chronology ensured consistency in the way the various countries' business cycle chronologies were determined and therefore enhanced the coherence of the international comparisons being undertaken. However, at the suggestion of an anonymous referee of an earlier version of the paper, the focus of the analysis was narrowed to Australia.

[5] The authors are grateful to an anonymous referee for pointing out this working paper to them.

[6] The consequences will vary, but, in certain applications, can be expected to be quite significant. For example, Pagan shows that one measure of the concordance between two countries' business cycles may be of the order of 80 per cent based on simply comparing the proportion of the time both countries are in expansions at the same time even if the two countries business cycles are actually unrelated. The apparently strong (but misleading) concordance is simply an artefact of the typically long durations of expansions of most countries.

Table I ECRI Business cycle chronologies for Australia

Classical cycle		*Growth cycle*	
Peak	*Trough*	*Peak*	*Trough*
June-51	Sep-52		Aug-52
Dec-55	Aug-56	May-54	Jul-56
Dec-60	Sep-61	Mar-57	Jan-58
Jun-74	Jan-75	Oct-59	Jun-61
Jun-82	May-83	May-62	May-63
Jun-90	Dec-91	Apr-64	Jan-66
		Feb-67	Jan-68
		Oct-68	Jan-72
		Oct-73	Jan-75
		Aug-76	Oct-77
		Dec-80	May-83
		Sep-85	Oct-86
		Jan-89	Apr-91
		Oct-94	Oct-96
		Mar-99	Nov-00

a negative coefficient which is consistent with expectations and indicates the possible presence of positive duration dependence; i.e. that the probability of a phase of the business cycle terminating increases as it becomes older. During expansions the estimated coefficient has a value of −0.0172 with a robust *t*-ratio of −1.79. Whist not significant at the conventional five per cent level of significance this result does provide some (albeit weak) evidence of the presence of positive duration dependency during periods of expansion.

The evidence supporting duration dependency is considerably stronger during periods of recession with the coefficient of the duration variable estimated at −0.2081 with a robust *t*-ratio –2.74. The much larger coefficient and its greater statistical significance provide quite strong evidence that duration exhibits a higher degree of both economic and statistical significance during periods of recession; i.e. that the influence of duration is asymmetric across the two phases of the classical business cycle. In fact the duration coefficient in recessions is approximately 12 times that in evidence in expansions. This makes intuitive sense given that the median length of expansions is around seven to eight years while the median length of recessions is around just nine months (refer to Table I). Thus, to the extent that expansions are duration dependent, one could expect the estimated dependence to be smaller than for recessions. On this basis, one would expect asymmetry in duration dependence between classical cycle expansions and recessions.

Comparison with Bodman (1998)

Whilst the results obtained above are broadly consistent with Bodman (1998), the divergence concerns the significance of the duration variable during periods of expansion. Bodman (1998) estimated a negative coefficient for duration, however, the estimated coefficient of −0.324 was only 1.49 times its robust standard error (of 0.17) and therefore provides little evidence that phase duration is statistically significant in explaining changes in the model's transition probabilities. To allow for a direct comparison with the results of Bodman (1998), the model is re-estimated using the same sample period studied by Bodman (1998). The results for the sample period, September 1959 to September 1997, are also outlined in Table II.

The results indicate that the data over this period, in comparison to the longer sample period initially used, display a higher degree of duration dependence both economically and statistically.

Table II Empirical results – classical cycle

	Whole data period		*Sept 1959–Sept 1997*	
	Model 1	*Model 2*	*Model 1*	*Model 2*
α_1	4.6858** (0.4496) [0.4705]	5.8686** (0.9992) [0.9481]	4.6250** (0.5099) [0.5175]	6.4094** (1.2243) [1.0682)
δ_1		−0.0172* (0.0105) [0.0096]		−0.0242** (0.0119) [0.0102]
α_2	2.2618** (0.4705) [0.4710]	3.9656** (1.1218) [0.8509]	2.3273** (0.5241) [0.5244]	4.1869** (1.3130) [1.0388]
δ_2		−0.2081** (0.1032) [0.0759]		−0.2129** (0.1143) [0.0873]
LL	−45.0128	−41.6091	−36.0176	−32.1336
λ		6.8047**		7.768**

Notes: Model 1: Constant transition probability model. Model 2: Time varying transition probability model using duration. * Indicates statistical significance at the 10% level of significance. ** Indicates statistical significance at the 5% level of significance. Asymptotic standard errors are in parenthesis and robust standard errors in square brackets.

The duration coefficients across both periods of the business cycle are still negative and are larger than those estimated for the initial sample period. For periods of expansion the duration coefficient is estimated to be −0.0242 with a robust *t*-ratio – which is now significant at the conventional levels of significance – of −2.37. The duration parameter during recessions is estimated at −0.2129, with a robust *t*-ratio of −2.45.

The significance of the duration coefficient during expansion, across both sample periods, is consistent with the results of Layton and Smith (2003) for the US. Layton and Smith also compared their results to the earlier work on the US of Durland and McCurdy (1994) who proposed the methodology adopted by Bodman (1998). Layton and Smith reconciled their results with Durland and McCurdy by suggesting that the different conclusions were the result of the increased precision obtained by investigating duration dependence within an established widely accepted business cycle chronology – *viz* the official NBER chronology for the US – in contrast to imperfectly inferring it from a single macroeconomic proxy. We speculate that the same argument may apply to the divergence in our results to those of Bodman (1998) regarding the significance of duration dependence in expansions.

Comparison with Cashin and Ouliaris (2004)

As noted in the introduction, Cashin and Ouliaris (2004) used non-parameteric tests applied to ex-post observable Australian classical business cycle phases (defined by applying the Bry-Boschan turning point algorithm to GDP to determine its chronology) and found that neither expansions nor recessions exhibited any significant duration dependence. This is a particularly unusual finding in relation to recessions and is at odds with the findings of many other earlier studies of different countries. In the case of Australia these results are also at odds with Bodman's (1998) finding of strong duration dependence in Australian recessions.[7] This is interesting given that both Bodman and Cashin and Ouliaris base their conclusions on phase switches in GDP alone – Bodman implicitly by virtue of his use of the Hamilton regime switching framework and Cashin and Ouliaris explicitly by their use of the Bry-Boschan GDP turning points. Our results, even though we use a different (and we would argue, preferable) basis for determining Australia's classical business cycle chronology, find strong evidence – as with Bodman for Australia, and as others have found for other economies – that recessions are characterised by a high degree of duration dependence.[8]

Influence of leading indicators

Explaining the transition probabilities based solely upon phase duration could possibly be problematic due to an omitted variable issue arising from ignoring the effects of other variables which could potentially explain business cycle phase changes. In addition to the duration of a phase, a more complete model might also include underlying economic fundamentals that might be believed to drive the business cycle. A number of economic indicators are believed to lead aggregate economic activity and hence provide an indication of the future phases of the business cycle. Such indicators are included within the modelling framework here through the use of a composite index which incorporates the information content of a number of different leading indicators.

Monthly observations of changes in the Australian long-leading index developed by ECRI are used to represent changes in such economic fundamentals driving future phase changes in the Australian business cycle. This index is commonly believed to anticipate changes in the business cycle by nine months and, as such, a nine-month moving sum of the month-to-month changes in the index is used to capture both the length and depth of swings in the index. Such a specification – by not separately including a large number of lags of the index in the model – economises on the use of parameters and reduces the potential problem of multicollinearity in the estimation.

The results of the model estimated solely with changes in the leading index (Model 1) and a combination of both phase duration and changes in the leading index (Model 2) are outlined in Table III. Considering model 1, the likelihood ratio statistic of 17.92 suggests that allowing the transition probabilities to vary over time according to changes in the leading index provides a superior explanation of the observed data in contrast to a constant transition probabilities model.

Importantly, considering model 2, phase duration continues to be a significant explanatory variable for phase changes during periods of recession even after changes in the leading index are included into the analysis. The magnitude of the phase duration coefficient increases to –0.3505 and remains statistically significant with a *t*-ratio of −2.02. In contrast, changes in the leading index appear to provide no significant informational content in explaining the transitional probabilities during recessions. Across both models 1 and 2 the coefficient of the leading index

[7] All the more surprising given that, even on the chronology used by Cashin and Ouliaris, Australia experienced six classical recessions over the period of their study, with the shortest three recessions each being two quarters and the others being four, five and six quarters in duration. Without formal statistical testing this would seem to be quite a strong clustering of recession durations around an average (median) duration of three quarters (two-thirds of the observations being within just one quarter of the average).

[8] Whilst not reported here in detail, an analysis of Canada, France, Germany and the US revealed evidence of recession duration dependence for all countries with some evidence of expansion duration dependence for the US and Germany. For Germany, the evidence for expansions was in fact more statistically significant than in the case of recessions. This is no doubt a reflection of the shorter expansions in that country in the post-war period.

Table III Classical cycle results – duration and long leading index

	Model 1	*Model 2*
α_1	5.2017**	5.2446**
	(0.7878)	(0.9691)
	[0.8063]	[0.8855]
δ_1		–0.0009
		(0.0119)
		[0.0073]
β_1	0.6882**	0.6794**
	(0.2061)	(0.2305)
	[0.1985]	[0.1907]
α_2	2.2568**	5.1330**
	(0.4831)	(1.6693)
	[0.4867]	[1.6402]
δ_2		–0.3505**
		(0.1749)
		[0.1731]
β_2	–0.0712	0.1623
	(0.0851)	(0.1548)
	[0.0708]	[0.1775]
LL	–36.0525	–33.6225
λ	17.92**	4.86*

Notes: Model 1: Time varying transition probability model using the leading index. Model 2: Time varying transition probability model using duration and the leading index. * Indicates statistical significance at the 10% level of significance. ** Indicates statistical significance at the 5% level of significance. Asymptotic standard errors are in parenthesis and robust standard errors in square brackets.

variable is smaller than its robust standard error. Additionally, when considering changes in the index together with phase duration, the estimated coefficient of the leading index variable is of the wrong sign. A negative coefficient is expected because an increase in the leading index, i.e. an improvement in economic fundamentals, would increase the likelihood of the economy switching into a period of expansion and hence decrease the probability of staying in a recession.[9] The results therefore would suggest that the transition probabilities during periods of recession are driven primarily by the length of time the recession has been in existence and hence exhibit a strong degree of positive duration dependence.

The leading index, however, exhibits strong predictive power in anticipating the end of expansions. Across both of the models estimated the coefficient is positive and statistically significant. The positive coefficient is expected because an increase in the leading index indicates an improvement in general business conditions and hence is expected to increase the probability of staying in an expansionary phase of the business cycle. The results for Model 2 indicate the predictive power of the leading index is sufficiently strong such that it mitigates any informational content

[9] The unexpected coefficient could possibly be a result of the high degree of multicollinearity experienced between duration and changes in the long leading index over periods of recession, evidenced by a correlation coefficient of 0.75.

that phase duration previously contained in explaining the transition probabilities. This suggests that the transition probabilities for expansionary phases of the business cycle are in fact driven by changes in the economic fundamentals of the economy, as captured by the leading index, rather than by phase duration.

An alternative interpretation of this result could be structured on the notion that changes in the leading index around turning points of the business cycle are a result of the aging process of the business cycle itself; i.e. it is some characteristic of lengthening duration that itself causes consequential changes in the leading index. An analogy here may be made with human mortality whereby, if one included a number of variables representing a person's general level of well being in a model describing the probability of dying, then it is conceivable that a person's age might itself be insignificant. This is because a person's general well being would contain all of the informational content of a person's age because it would be expected that a person's general wellbeing is significantly correlated with their age. In the context of the business cycle then it might be argued that the changes in the leading index are a reflection of the aging process of the business cycle itself!

Another point is also worthy of mention here. In investigating the issue of duration, one could argue the key question is simply whether phase durations cluster around some average duration such that the longer the duration of the business cycle phase the greater is its probability of terminating. This would suggest that, irrespective of whether the underlying cause of an observed phase switch is actually the result of some changing economic fundamentals, the key issue is simply whether or not the evidence on phase switches supports the existence of a systematic relationship between the probability of experiencing a phase switch and phase duration itself.

On this argument, whilst the leading index may well display a significant ability to explain the transition probabilities of the phases of the business cycle, explaining – and forecasting – business cycle phase shifts is not the essential research question being studied as far as duration dependency is concerned. Instead, it is simply whether the phases of the business cycle are characterised by duration dependence and, as such, the insignificance of the duration variable when considered with the leading index does not necessarily nullify the previous evidence found regarding positive duration dependence in expansions. On this argument, a reasonable conclusion then to reach from the model incorporating both duration and the leading index is simply that duration is evidently considerably less important for expansions than it is for recessions.

b) The growth cycle

Results for the Australian growth cycle – incorporating duration alone – are provided in Table IV (Model 2). The inclusion of duration within the model significantly improves the explanatory power of the model as recognised by the likelihood ratio statistic of 19.997. The coefficients of the duration variable across both phases of the growth cycle are negative and strongly significant. During fast growth phases the estimated duration coefficient increases from the marginal impact evidenced within the classical cycle to have a much more significant impact with a coefficient of -0.0763 and a corresponding *t*-ratio of 3.24. Duration remains significant during slow growth phases with a coefficient of -0.0994 and a robust *t*-ratio of 4.19.

Comparison with Cashin and Ouliaris (2004)

The results regarding duration dependence within the growth cycle presented here, we believe, have more intuitive appeal than the previous results reported for the Australian growth cycle by Cashin and Ouliaris (2004). Cashin and Ouliaris (2004) suggest that only slow growth phases

Table IV Growth cycle results

	Model 1	*Model 2*
α_1	3.0812**	4.3814**
	(0.2711)	(0.5984)
	[0.2688]	[0.5564]
δ_1		−0.0763**
		(0.0257)
		[0.0235]
α_2	2.9482**	4.4984**
	(0.2736)	(0.6210)
	[0.2731]	[0.5122]
δ_2		−0.0994**
		(0.0291)
		[0.0237]
LL	−113.0895	−103.0909
λ		19.9972**

Notes: Model 1: Constant transition probability model. Model 2: Time varying transition probability model using duration. * Indicates statistical significance at the 10% level of significance. ** Indicates statistical significance at the 5% level of significance. Asymptotic standard errors are in parenthesis and robust standard errors in square brackets.

of the growth cycle exhibit duration dependence, and that it is negative duration dependence; i.e. the longer the slow growth phase, the more likely it is that the phase will continue! Intuitively, however, one would expect the presence of positive duration dependence within the growth cycle because the fast growth and slow growth phases of the growth cycle are the result of de-trending the relevant economic time series involved in measuring the business cycle. As such, by construction, one would expect the business cycle measure to revert to the average level of growth in due course rather than the other way around. The results presented here provide further support to the conclusions of Abderrezak (1998).

Symmetry or asymmetry of duration dependence in the Australian growth cycle

As noted earlier, the impact of duration appears to be strongly asymmetric across the two phases of the classical business cycle. This does not seem to be the case for the Australian growth cycle.

For the growth cycle, as has been previously mentioned, whilst it is expected that the phases of the growth cycle would exhibit positive duration dependence as a result of the de-trending procedures applied to the data, de-trending, *per se*, does not necessarily imply duration dependence symmetry. Duration dependence symmetry would require that the speed with which economic growth returns to trend from above would need to be approximately the same as from below. This will depend upon the pattern of growth around trend in evidence in the historical record and the similarity of speed of return to trend from above and below trend may or may not be in evidence. Nonetheless, the results presented here do also suggest that, unlike the classical cycle, there is evidence in favour of duration dependence symmetry for the growth cycle; i.e. the difference between the coefficients of duration across the two phases is not statistically significant. Application of the likelihood ratio test of the restriction that the coefficients are equal results in a likelihood ratio statistic of 1.4544 which, with a critical value of 3.841 (2.706) at the five per cent (ten per cent) level of significance, suggests there is insignificant evidence to reject the restriction.[10]

Thus, for Australia, once the long-term trend has been removed, the average length of below trend periods of growth is apparently about the same as the periods of above trend growth. In fact, of the 15 growth accelerations in the post-war historical record to date, the median length is 21 months, and, of the 14 growth slowdowns the median length is 20 months (again, refer Table I). Thus, the closeness of the values of the duration coefficients for the growth cycle – as well as the similarity in statistical significance of the coefficients – is not too surprising.

IV. Conclusions

This paper re-examines duration dependency within the phases of the Australian business cycle (both classical and growth cycles) using a multivariate logit regime switching approach. Utilising the business cycle chronologies compiled by ECRI for the post-WW2 period, significant evidence is found suggesting the presence of duration dependency. Importantly, in the case of Australia's classical cycle, this contrasts quite strongly with the recent finding of the absence of duration dependence – in either expansion or recession phases – by Cashin and Ouliaris (2004). In the case of the growth cycle we find strong evidence of duration dependence, the effect of which is symmetric across both phases of the cycle.

A number of specific conclusions emerge from the empirical results presented in this paper.

Firstly, the recessionary phases of classical business cycles are characterised by strong positive duration dependency. The evidence on the recessionary phase of the Australian classical business cycle supports the characterisation that the probability of a recession terminating increases the longer the recession is in existence. As far as the expansionary phase is concerned the evidence is more muted. Interestingly, using the same sample period as that of Bodman (1998), but in contrast to him, apparently statistically significant evidence is found for duration dependence in Australia's expansions. Using a longer sample period, the evidence for Australian expansions is somewhat weaker but still appears significant at the ten per cent level.[11]

Secondly, the leading index employed here exhibits significant power in predicting the termination of expansionary phases of the business cycle. The leading index is a sufficiently powerful driver of the transition probability for expansions that any informational content which may be present in current phase duration alone is substantially mitigated. The leading index, however, does not have any informational content for the probability of the termination of a recession beyond that contained in the current duration of the recession.

Thirdly, the Australian growth cycle exhibits significant positive duration dependence across both phases of the cycle. Again, the results contradict the recent results of Cashin and Ouliaris (2004) who found no duration dependence for expansions and *negative* duration dependence for recessions! We would suggest that the current results seem to accord more with intuition given that the growth cycle chronology is defined in terms of detrended fluctuations in the business

[10] The growth cycles of France, Canada, Germany and the US also all displayed evidence of strong positive duration dependence for both expansions and recessions with the null of duration dependence symmetry accepted for Germany and the US but rejected for France and Canada.

[11] The unreported evidence from the analysis of the economies of the US, Canada, France and Germany was very similar to Australia for recessions. There was also some evidence of duration dependence in expansions for the US, and some quite strong evidence of this for Germany.

cycle measure. Our results are, however, consistent with those of Abderrezak (1998). We also found that duration dependence seems symmetric across both phases of the growth cycle.[12]

A natural question to ask is why these results might differ from earlier Australian studies of this issue (*viz* Bodman, and Cashin and Ouliaris). After all, at one level one could simply argue that a different methodology using a different chronology might well lead to different conclusions. This may well be so. For example, using different chronologies to those used here may well yield an estimated model with different statistical properties and possibly different qualitative conclusions. However, we believe the current results have a certain amount of intuitive appeal and, by virtue of incorporating the influence of the informational content in a leading indicator into the analysis, represents a more complete analysis of the issue of duration dependence in the Australian business cycle. The resulting estimated model potentially may therefore provide a richer model of the likelihood of imminent turning points in the cycle.

From the standpoint of gaining a better economic understanding of the nature of Australia's business cycle, an important conclusion of the paper then is that, once information from leading economic indicators is taken into account in assessing the likelihood of an imminent end to an expansion, the current duration of the expansion is of little to no informational value. On the other hand, once the Australian economy is in a recession, it would seem the best predictor of its end is very much its current length.

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[12] The evidence in support of positive duration dependence in the growth cycle was also strong across all of the other countries studied. However, only the US and Germany displayed duration dependence symmetry in their growth cycles.

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