SHOCK PERSISTENCE IN AUSTRALIAN OUTPUT AND CONSUMPTION

by
Luigi Ermini

No. 136  March 1990

DEPARTMENT OF ECONOMICS
SHOCK PERSISTENCE IN AUSTRALIAN OUTPUT AND CONSUMPTION

by

Luigi Ermini

No. 136 March 1990

Abstract

The distinction between transitory and permanent shocks bears important implications for economic analysis in general, and for forecasting and policy issues in particular. This distinction is at the centre of the debate on which class of models is best suited to represent economic variables: stationary models around a deterministic trend, or stationary models around a stochastic trend. The debate is here focussed on the Australian case. It is found that both output and consumption are characterized by stochastic trends, but without transitory component. This corresponds to a measure of persistence equal to one for both variables.
## CONTENTS

1. Introduction
2. Measures of Persistence
   2.1 Stationary Processes
   2.2 Integrated Processes
   2.3 Unobserved Component (UC) Processes
   2.4 Other Measures of Persistence
3. The Empirical Results
4. Conclusions
   References
   Addendum
1. Introduction

Within the modern view that economies are "driven" by unanticipated shocks, or innovations, an important question is to establish whether the effects of such shocks fade away after a reasonably long period of time (transitory shocks), or persist throughout the future life of the economy (persistent shocks). The distinction is clearly relevant for economic analysis, particularly for forecasting and policy issues. A transitory shock will affect only the short-run forecast of the relevant variable; a persistent shock will affect the long-run forecast as well. Similarly, if the effect is transitory, a shock-generating policy - whether fiscal or monetary - will only have a short-term benefit; otherwise, it will affect the economy permanently.

The distinction between transitory and permanent shocks provides the econometrician with a richer range of models available to represent economic variables. Based on the conventional practice of treating economic output as decomposable into a permanent trend component (the growth component) and a transitory stationary component (the business cycle), this distinction is in fact at the center of the current debate on whether economic variables in general are best represented by stationary processes around a deterministic trend or by stationary processes around a stochastic trend. In the former case, the shocks are transitory, and affect only the transitory component of the variable (the stationary process). Within this view, for example for the case of output, shock-generating policies would be only policies of output stabilization around the natural rate of growth. In the latter case, the shocks are permanent, and affect in various degrees both the transitory and the permanent component of the relevant variable. Again for the case of output, within this view a shock-generating policy would affect the business cycle and simultaneously alter the rate of growth of the economy. Not surprisingly, the debate about which class of models is best suited to represent economic variables is far from settled. For a review, see Watson [1986], Clark [1987], King, Plosser, Stock and Watson [1987], Campbell and Mankiw [1987a, 1987b], Cochrane [1988], and Stock and Watson [1988].

The purpose of this paper is to focus the debate on the Australian case, by concentrating on the measure of shock persistence in disposable income and private consumption. Studies on the dynamic properties of Australian income and consumption
are surprisingly rare. For the specific case of consumption, the few reported studies have concentrated on the question of whether the Australian consumption follows the permanent income hypothesis or not (Johnson [1983], McKibbin and Richards [1988]), without attempting to understand if in fact consumption is characterized by a more complex dynamic structure. The paper is organized as follows: Section 2 reviews the theoretical aspects associated with the notion of shock persistence, for three classes of stochastic processes (stationary, integrated, and unobserved component (UC)). Section 3 reports the empirical results, including a test for cointegration. Section 4 provides some concluding remarks.

2. Measures of persistence

2.1. Stationary processes

Consider the Wold's decomposition of a wide-sense zero-mean stationary process \( \{y_t\} \):

\[ y_t = \sum_{j=1}^\infty c_j \varepsilon_{t-j}, \quad c_0 = 1, \tag{1} \]

where \( \varepsilon_t \) is a zero-mean white noise series of variance \( \sigma^2; \varepsilon_t \) is termed the "innovation" or shock in \( y_t \), as it corresponds to that part of \( y_t \) that could not be anticipated on the basis of the (univariate) information set available at time \( t-1 \), \( \varepsilon_t = \varepsilon_t - E(\varepsilon_t | I_{t-1}) \), with \( E \) the expectation operator. A compact notation for (1) is \( y_t = C(\varepsilon) \varepsilon_t \), where \( C(\varepsilon) = \sum_{j=0}^\infty c_j \varepsilon_t^j \) is a polynomial of infinite order in the backward operator \( \varepsilon \) (i.e., \( \varepsilon^2 \varepsilon_t = \varepsilon_{t-1} \)). Stationarity requires \( \Sigma c_j < \infty \).

The Wold's decomposition essentially states that the value of the variable \( y_t \) at any period \( t \) is made up of a linear combination of all past shocks to the series. Thus, the past shock \( \varepsilon_{t-j} \) contributes the amount \( c_j \varepsilon_{t-j} \) to the current value of the variable. Since stationarity implies (Harvey [1981], Granger and Newbold [1986]):

\[ \lim_{t \rightarrow \infty} c_j = 0, \tag{2} \]

it clearly follows that for a wide-sense stationary process the effect of shocks fades away with time. Thus, stationarity implies transitory shocks, or, equivalently, implies zero persistence. Occasionally, this concept is also expressed by saying that a wide-sense stationary process has "finite memory": the very distant past does not affect the present. Alternatively, this same concept is also expressed as "mean reversion": if from time \( t \) on, \( y_t \) is not subject to further shocks, its value at time \( \tau + k \) becomes \( y_{\tau+k} = \sum_{j=1}^\infty c_j \varepsilon_{\tau-j} \), from which \( \lim_{\tau \rightarrow \infty} y_{\tau+k} = 0 \). Thus, a stationary process \( y_t \) tends with time to revert towards its mean value (zero).

For an interesting implication of transitory shocks, consider the optimal mean-square (univariate) forecast of \( y_{\tau+k} \), based on the univariate information set \( I_t \) (Granger and Newbold [1986]):

\[ \hat{f}_{\tau+k} = \sum_{j=0}^\infty c_j \varepsilon_{t-j}. \]

From (2) it follows that

\[ \lim_{\tau \rightarrow \infty} \hat{f}_{\tau+k} = 0, \]

i.e., the long-run optimal forecast coincides with the unconditional expectation of the variable. Since under mean-square rule the optimal forecast is in fact the conditional expectation, stationary processes exhibit the property that their conditional mean ultimately reverts to their unconditional mean. An identical result is obtained when the unconditional mean is, more generally, time-varying according to a (necessarily) deterministic function of time, \( T_t \). In this case, \( y_t = T_t + C(\varepsilon) \varepsilon_t \), and the long-run conditional mean again reverts towards the unconditional mean.

This property defines the principle on which the decomposition of an economic variable into permanent and transitory components is based: the permanent component coincides with the long-run forecast; the transitory component with the "detrended" short-run forecast. In our case:

\[ y_t = y^p_t + y^T_t, \tag{3} \]

with

\[ y^p_t = T_t \quad (\text{permanent component}) \]

\[ y^T_t = C(\varepsilon) \varepsilon_t \quad (\text{transitory component}). \tag{4} \]

Note that in the present case of stationary processes, by construction the permanent component \( y^p_t \) is deterministic, and hence unaffected by the shock \( \varepsilon_t \). As anticipated, this decomposition is often called "stationarity around a deterministic trend".

Macrovariables such as income and consumption typically exhibit an upward trend. Within this class of models, then, their permanent component would be specified as a monotonically increasing function of time. Standard practice for GDP, for example, is to regress its logarithm against a linear time trend, to capture the
natural rate of growth. The detrended series then - provided it is stationary - is identified with the transitory component, and, for the case of output, is interpreted as describing the business cycle. However, as indicated in recent literature (particularly, Durlauf and Phillips [1988]), this practice can lead to serious specification and estimation problems.

2.2. Integrated processes

A natural extension of the previous class of stationary processes around a deterministic trend is to hypothesize that the trend or permanent component is also a stochastic process. Central to these stochastic-trend models is the class of "integrated processes", whose non-stationarities are eliminated by data differencing rather than by data detrending. A process \( y_t \) is said to be integrated of order \( d \), \( I(d) \), if it requires \( d \) differencings in order to become wide-sense stationary. That is, if

\[
(1-B)^d y_t = m + C(B) \epsilon_t,
\]

with \( \epsilon_t \) a zero-mean white noise, \( m \) a constant and \( C(B) \) a polynomial (possibly of infinite order) in the backward operator \( B \). Conventionally, a stationary process is also indicated as \( I(0) \), as it does not require any differencing to become stationary (\( d=0 \)). As almost all relevant macrovariables are found to be typically \( I(1) \) (for example, Nelson and Plosser [1982]), in what follows only \( I(1) \) processes will be considered. That is, processes whose first differences \( \Delta y_t = (1-B) y_t \) are stationary:

\[
(1-B) y_t = m + C(B) \epsilon_t. \quad (6)
\]

Again, \( \epsilon_t \) is the innovation in the value of \( y_t \), i.e. that part of \( y_t \) that could not be predicted on the basis of the previous (univariate) information set \( I_{t-1} = \{\epsilon_{t-1}, j \geq 1\} \); that is, \( \epsilon_t = y_t - E(y_t | I_{t-1}) \). Note that this conditional expectation is equal to \( y_{t-1} + \sum \epsilon_{t-i} \), but \( y_{t-1} \) is a linear combination of known past shocks and hence is indirectly part of \( I_{t-1} \) (see also (8) below).

\( I(1) \) processes exhibit a set of completely different properties from stationary processes. For instance, their unconditional variance is a linearly increasing function of time, with slope \( \sigma^2 \); their unconditional mean is a linear function of time with slope \( m \) (see (7) and (8) below). In particular, regarding shock persistence, it will be seen that the shocks to \( I(1) \) processes persist forever, as opposed to \( I(0) \) processes whose shocks are transitory. To illustrate, consider the simple case of a random walk with drift, defined for \( C(B) = 1 \):

\[
y_t - y_{t-1} = m + \epsilon_t.
\]

On solving this difference equation, with zero initial value without loss of generality, we get

\[
y_t = m \sum_{j=0}^{t-1} \epsilon_j.
\]

Thus, in addition to a linear time trend, the value of the variable at any period is made up of the undiscounted accumulation of all the past shocks. Each shock \( \epsilon_{t-j} \) contributes its full value, and not, as in stationary processes, its discounted value \( C(B) \epsilon_{t-j} \). As they are not discounted by the coefficient \( \epsilon \), the shocks to a random walk persist forever. Further, as the fraction of the value of the variable at any time is equal to one (and is constant throughout the future life of the process), one is an obvious choice as the measure of persistence for random walk shocks.

For the general \( I(1) \) process (6) we have:

\[
y_t = m \sum_{j=0}^{t-1} \epsilon_j + \sum_{j=0}^{t-1} \epsilon_{t-j}.
\]

The corresponding shock effect is now equal to \( \sum \epsilon_j \) for the shock \( \epsilon_{t-j} \), and the ultimate or long-run persistence is equal to \( C(1) \). Thus \( C(1) \) can be taken as an obvious measure of persistence for any class of models. In particular, for a random walk, since by construction \( C(B) = 1 \), \( C(1) = 1 \) yields the persistence of one as outlined above. For a stationary process, it is known (see, for example, Granger and Newbold [1986]) that for overdifferenced stationary processes - that is, for stationary processes whose first differences are represented as in (6) - \( C(1) = 0 \); this yields the zero measure of persistence for stationary processes, as outlined in the previous section. For a general \( I(1) \) process, \( C(1) \) could be greater or less than one. For example, US monthly consumption is well fitted by the (univariate) model \( \Delta c_t = \epsilon_t - 0.21 \Delta c_{t-1} \) (Ermini [1989a]), so that its measure of persistence is equal to 0.789. This entails that every dollar of monthly unexpected consumption expenditure signals a permanent monthly increase of 79 cents in consumption expenditures for the entire remaining life.

As shocks to integrated processes persist forever, integrated processes are not mean reverting (but their differences are). Further, as their optimal (mean-square, univariate) forecast is
The permanent component is always a random walk (possibly with drift); the transitory component is a stationary process, whose effects on the variable decay in the long-run. Notice also that the permanent component captures, through \( C(t) \), the shock persistence of the overall process \( y \). To distinguish from the case of the previous section, this class of models is referred to as "stationarity around a stochastic trend".

### 2.3. Unobserved component (UC) processes

In the decomposition (14) of the previous section, both permanent and transitory components are "driven" by the same innovation \( \varepsilon \). A natural extension of the stochastic-trend approach is to introduce the possibility of two different innovations driving the two components:

\[
(1-B)y_t = m + \eta_t,
\]

\[
y_t^* = C^*(B) \eta_t,
\]

with \( \nu_t \) and \( \eta_t \) exhibiting in general various degrees of correlation. (This class of models derives its name from both \( \nu_t \) and \( \eta_t \) being unobservable, as opposed to \( \varepsilon_t \) in (14).) The relationship between the shock \( \varepsilon_t \) to the overall process \( y_t \) and the two component-specific shocks is given by

\[
C(B) \varepsilon_t = \nu_t + (1-B) C^*(B) \eta_t.
\]

The case of perfect correlation between \( \nu_t \) and \( \eta_t \) is readily seen to correspond to the Beveridge-Nelson decomposition (14) of the previous section. The opposite case of no correlation - which is the usual case of UC models considered in the literature - corresponds to a restricted version of the general model (6). The condition \( \varepsilon \nu' \eta = 0 \) in fact is shown in Watson [1986] to place the restriction that the spectral density of \( \Delta y \) reaches a global minimum at the zero frequency. Hence, under this condition of uncorrelatedness, not all polynomials \( C(B) \) are suited to represent \( y_t \) as in (6).

With regard to persistence, Campbell and Mankiw [1987] show that for UC models with uncorrelated components the measure of persistence \( C(0) \) is at most equal to one.

Under the condition of uncorrelatedness, the identity (16) can be used to calculate the polynomial \( C(B) \) from the knowledge of \( C^*(B) \), by applying Granger's Lemma (Granger and Newbold [1986], p. 29), and by solving for the implied moving average coefficients. Correspondingly, knowing \( C(B) \), the measure of persistence \( C(0) \) for the UC model with uncorrelated components can be evaluated. For example, consider the
model, often assumed in the literature, that output follows a random walk corrupted by an uncorrelated white noise measurement error (for example, Sargent [1979]). In the notation of this section, this corresponds to:

\[(1-\theta)y_t^{\prime} = \nu_t \]
\[y_t^{\prime} = \eta_t \]
\[\varepsilon y_t, \eta_t = 0. \]

From (16) it follows that

\[C(\theta) = \sigma_{\varepsilon}^2 - \nu_t + (1-\theta) \eta_t = (1+\theta) \eta_t, \]

that is, observed first differences of output follow an IMA(1,1) process, with autocovariances

\[R_{y_0}(0) = (1+\sigma_{\varepsilon}^2) \sigma_{\varepsilon}^2 = \sigma_{\varepsilon}^2 + 2 \sigma_{\eta}^2 \]
\[R_{y_1}(1) = \sigma_{\varepsilon}^2 \sigma_{\eta} = -\sigma_{\eta}^2. \]

For \(c < 0\) (implied by the first-lag autocovariance), the solution to the equation

\[(1+\sigma_{\varepsilon}^2) \sigma_{\eta}^2 = \psi (\sigma_{\eta}^2, 2 \sigma_{\eta}^2) = 0 \]

is

\[c = -1 + \frac{1}{2} \frac{1}{\sigma_{\varepsilon}^2} \frac{\sigma_{\eta}^2}{\sigma_{\eta}^2}, \]

where \(\psi = \sigma_{\varepsilon}^2 \sigma_{\eta}^2\) is the signal-to-noise ratio. For \(\psi > 0\), (20) implies \(-1 \leq c \leq 0\). Correspondingly,

\[C(\psi) = 1 + c = \sqrt{\psi} (\sqrt{\psi} + \frac{1}{2} \sqrt{\psi}) \frac{1}{\sqrt{\psi}} \]

which varies between 0 and 1. The case \(C(\psi) = 0\) corresponds to \(\sigma_{\varepsilon}^2 = 0\) (equivalent to \(\theta = 0\), i.e. output is just white noise, \(y = \eta_t\)); the case \(C(\psi) = 1\) corresponds to \(\sigma_{\varepsilon}^2 = 0\) (equivalent to \(\eta_t = 0\), i.e. output is an uncorrupted random walk, \(\Delta y = \nu_t\)).

The range of possible models can be further expanded by introducing higher-order integrated processes, \(I(d)\) with \(d > 1\). For example, the model presented in Clark [1987], in which the drift of the permanent component in (14) or (15) also follows a random walk, corresponds to an \(I(2)\) integrated process. However, as output and consumption clearly appear to be \(I(1)\) and not \(I(2)\) (see Section 3), this expanded class of models will not be considered in this paper.

2.4. Other measures of persistence

The calculation of the measure of shock persistence \(C(\psi)\) for an UC model with uncorrelated components, as outlined in the previous section, could be quite a cumbersome task. On the other hand, if one wants to compare measures of persistence across different classes of models - for example between an UC model with uncorrelated components and an unrestricted \(I(1)\) - a direct comparison is meaningless: the persistence of the UC model is given by the persistence of its permanent component, which is related to the shock \(\nu_t\), while the persistence of the unrestricted \(I(1)\) is related to the different shock \(\eta_t\). To overcome this problem, another measure of shock persistence, introduced by Cochrane [1988], can be used. Cochrane's measure is defined as

\[V = \lim_{\psi \to 0} \frac{1}{\psi + \frac{1}{\psi}} \frac{\nu_t^2 + \eta_t^2}{\nu_t^2 + \eta_t^2}. \]

It can be shown (see also Campbell and Mankiw [1987b]) that this measure is in fact equivalent to the explained variation of \(\Delta y\), due to its permanent component:

\[V = \frac{\text{var} \Delta y_{\text{fix}}}{\text{var} \Delta y}. \]

Under the Beveridge-Nelson decomposition (14), \(\Delta y_{\text{fix}} = C(\psi) \varepsilon_t\), so that \(\text{var} \Delta y_{\text{fix}} = C^2(\psi) \varepsilon_t\). By letting \(1 - R^2 = \sigma_{\varepsilon}^2 / \text{var} \Delta y\) be the unexplained variation of \(\Delta y_t\), we get

\[V = C^2(\psi) (1 - R^2). \]

In the case of UC models with uncorrelated components, \(\text{var} \Delta y_{\text{fix}} = \sigma_{\varepsilon}^2\), so that

\[V = \frac{\sigma_{\varepsilon}^2}{\text{var} \Delta y}. \]

It can be shown that \(\sigma_{\varepsilon}^2 = C^2(\psi) \varepsilon_t\), where \(C(\psi)\) corresponds now to the polynomial \(C(\theta)\) of the identity (16) 1. It follows that instead of comparing persistence across models with the measure \(C(\psi)\), persistence can be compared with the measure \(V\), obtained as in (23) for unrestricted \(I(1)\) models, and as in (24) for UC models with uncorrelated components.

3. The empirical results

The first stage of the empirical work is to check whether output and consumption belong to the class of stationary models around a deterministic trend or to the class of stationary models around a stochastic trend. This check is carried out through the following steps: (a) regress separately \(\log y\) and \(\log c\) against a linear time trend, and check the stationarity of the residuals 2. More precisely, the residuals are subject to a

---

1 In place of a formal proof, consider again the example described in (17). We have:
\[C^2(\psi) \varepsilon_t = (1+\psi)^2 \varepsilon_t = (1+\psi)^2 \varepsilon_t = 2 \sigma_{\varepsilon}^2 \varepsilon_t. \] But from (19), \((1+\psi)^2 \varepsilon_t = \sigma_{\varepsilon}^2 + 2 \sigma_{\varepsilon}^2 \varepsilon_t + \sigma_{\varepsilon}^2 \varepsilon_t\)

2 The choice of logarithmic values - common in the literature - is based on the observation
unit root test to check for first-order integration; (b) apply to both \( (y_t, c_t) \) and \( (\log y_t, \log c_t) \) a unit root test without detrending, to check directly for first-order integration. This double check enables us to confirm that both output and consumption are integrated processes, and hence that their regression against a linear trend is not spurious in the sense of Durbin and Phillips [1988], but indeed it amounts to detrending integrated processes.

As both output and consumption (and their log values) appear to be \( I(1) \), that is to belong to the class of stochastic-trend models, the second stage is to check whether the two series are best represented as unrestricted \( I(1) \) models, or as UC models with uncorrelated components. Since the best (mean-square, univariate) unrestricted model for both \( (y_t, c_t) \) and \( (\log y_t, \log c_t) \) turns out to be a random walk with drift, the estimation as UC models becomes redundant. Both output and consumption have a clear measure of persistence equal to one, with apparently no transitory component around their stochastic trend. Because of this finding, equally redundant are in our case the non-parametric estimations of \( C(1) \) (or \( V \)), of the type proposed by Campbell and Mankiw [1987b].

Finally, the third stage is to establish whether output and consumption have a common stochastic trend. This is done by simply testing output and consumption for cointegration. The empirical results are:

(a) *detrending logarithmic values:* for output the OLS estimates of the log-linear model are (r-statistics in parenthesis):

\[
\log y_t = 9.95 + 0.995 t, \quad (R^2 = 0.99, D.W. = 0.032, \text{r-statistic} = 62.5) \]

(25)

with \( R^2 = 0.97, D.W. = 0.052 \), and the first-order autocorrelation of the residuals \( \rho = 0.965 \). These values clearly indicate the presence of a unit root in the residuals. In fact, a formal augmented Dickey-Fuller [1981] test confirms the case (Table 1). Considering

\[
\Delta y_t = \alpha_0 + \alpha_1 t + \alpha_2 y_{t-1} + \sum_{j=1}^{\infty} \beta_j \Delta y_{t-j},
\]

(26)

where \( y_t \) is the residual from (25), and testing \( H_0: \alpha_1 = \alpha_2 = 0 \) (i.e., unit root with zero intercept) against \( H_1: (\alpha_1, \alpha_2) \neq 0 \), yields a practically zero value for the test statistic \( \varphi \). As the critical value at the 1% level is 6.50 (Dickey and Fuller, p. 1063, for sample size 100), the test strongly accepts the hypothesis that the residuals have a unit root and a zero intercept. The same result is obtained with a simple Dickey-Fuller test.

For consumption the estimated log-linear model is

\[
\log c_t = 9.43 + 0.0995 t, \quad (R^2 = 0.98, D.W. = 0.034, \text{r-statistic} = 75.9) \]

(27)

with \( R^2 = 0.98, D.W. = 0.034, \text{and } \rho = 0.970 \). Again, the results reported in Table 1 confirm that the residuals have unit root with zero intercept.

(b) *unit root tests on \( (y_t, c_t) \) and \( (\log y_t, \log c_t) \):* these are reported in Table 1 as well. In this case, because of a significant intercept, the tests are based on the null \( H_0: \alpha_1 = \alpha_2 = 0 \) against the alternative \( H_1: (\alpha_1, \alpha_2) \neq 0 \). All test results strongly accept the null hypothesis of unit root with significant intercept for the four listed series.

(c) *model selection:* Table 2 report the results of the likelihood ratio tests for \( \log y_t \). The null hypothesis of random walk is never rejected against any other richer representation, even without penalizing the latter models for the number of parameters. (Note the plausible existence of a common root in the ARIMA(1,1,1) model). Similar results, available from the author, apply for \( c_t, y_t \) and \( \log c_t \). So, the evidence strongly supports that Australian output and consumption are random walks with drift. Precisely:

\[
\lambda y_t = 366 + \varepsilon_t, \quad (\text{residuals } \hat{\varepsilon} = 0.016) \]

(28)

\[
\Delta \log y_t = 0.0098 + \varepsilon_t, \quad (\text{residuals } \hat{\varepsilon} = 0.013) \]

\[
\lambda c_t = 205 + \varepsilon_t, \quad (\text{residuals } \hat{\varepsilon} = 0.016) \]

\[
\Delta \log c_t = 0.0025 + \varepsilon_t, \quad (\text{residuals } \hat{\varepsilon} = 0.0085)
\]

Note that, in line with the life-cycle hypothesis of consumption smoothing, Australian consumption indeed appears to have lower volatility than income.

These results indicate that both income and consumption do not have transitory components, but strictly comprise a permanent component only. For this reason, the alternative estimation of the corresponding UC models with uncorrelated components is considered redundant (the restriction of global minimum at the zero frequency is satisfied by the spectral density of white noise). Incidentally, the apparent lack of business cycle in Australian GDP - as resulting from (28) - warrants further investigation.
on the nature and origin of such a random walk. For example, recalling the model described in (17), how can we reconcile a random walk with the almost absolute certainty that income data are corrupted by measurement errors?

(a) cointegration: the estimated cointegrating equation is

$$c_t = 181 + 0.59 y_t + u_t$$  

(30)

with $D.W. = 0.247$. The low value of the Durbin-Watson statistics indicates that the residual $u_t$ has a unit root at the 1% and 5% levels. (cf. the critical values in Table II and III of Engle and Granger [1987]). Only at the 10% level can we accept cointegration. The same negative conclusion of non-cointegration is obtained with a formal augmented Dickey-Fuller test, as described in Engle and Granger. The $t$-statistics of the coefficient $\phi$ in $\Delta u_t = \phi u_t + \sum_{j=1}^{4} \beta_j \Delta u_{t-j}$ is -1.22, well below the critical values of either tables II and III of Engle and Granger up to the 10% level. Note, incidentally, that with a linear coefficient between $c_t$ and $y_t$ of 0.59, the corresponding linear combination of the two drifts for $c_t$ and $y_t$ is 10, i.e. the residuals are $t(1)$ with a small drift.

Non-cointegration between Australian output and consumption was also found by McKibbin and Richards [1988], even if they use a different set of data. Rejection of cointegration entails that output and consumption do not have a common stochastic trend, and also that no error-correction mechanism exists, based on the deviations between these two variables, to push both of them back towards equilibrium. Finally, as discussed in Ermini [1989], rejection of cointegration entails rejection of the permanent income hypothesis. This can be easily seen by considering that, as income follows a random walk, it coincides with permanent income; then, rejection of cointegration is equivalent to rejection of a linear relation between consumption and permanent income.

4. Conclusions

It is found that Australian domestic output and private consumption are not stationary processes around a deterministic trend. In particular, this finding rejects the conventional identification of the business cycle with detrended output. Both processes are unambiguously found to be integrated of order one; therefore they belong to the class of stationary processes around a stochastic trend (or stochastic permanent component). However, it is found that the best univariate representation for both series is the random walk model with drift. This implies that output and consumption have no transitory component, and that their measure of shock persistence is one. Particularly, with regard to output, the lack of a transitory component would surprisingly indicate the absence of business cycles in the Australian economy. Finally, it is found that output and consumption are not cointegrated. This implies rejection of the permanent income hypothesis for the Australian case, and rejection of the possibility that output and consumption share a common stochastic trend.
<table>
<thead>
<tr>
<th>SERIES</th>
<th>RESIDUALS FROM (25)</th>
<th>RESIDUALS FROM (27)</th>
<th>$y_t$</th>
<th>$\log y_t$</th>
<th>$c_t$</th>
<th>$\log c_t$</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$H_0$</td>
<td>$H_A$</td>
<td>$H_0$</td>
<td>$H_A$</td>
<td>$H_0$</td>
<td>$H_A$</td>
</tr>
<tr>
<td>$a_0$</td>
<td></td>
<td>-0.001378 (0.36)</td>
<td>-</td>
<td>0.00208 (1.61)</td>
<td>365.12 (9.21)</td>
<td>0.0098 (9.56)</td>
</tr>
<tr>
<td>$a_1$</td>
<td></td>
<td>-0.000127 (0.15)</td>
<td>-</td>
<td>-0.00004 (-1.88)</td>
<td>15.22 (1.04)</td>
<td>-</td>
</tr>
<tr>
<td>$a_2$</td>
<td></td>
<td>-0.0267 (-1.32)</td>
<td>-</td>
<td>-0.0187 (-0.97)</td>
<td>-</td>
<td>-0.0346 (-1.92)</td>
</tr>
<tr>
<td>$B_1$</td>
<td></td>
<td>-0.0372 (-0.41)</td>
<td>-0.028 (-0.30)</td>
<td>-0.1728 (1.81)</td>
<td>1.525 (-0.96)</td>
<td>-0.082 (-0.80)</td>
</tr>
<tr>
<td>$B_2$</td>
<td></td>
<td>-0.042 (0.46)</td>
<td>-0.0534 (0.58)</td>
<td>-1.295 (-1.34)</td>
<td>-1.469 (-1.46)</td>
<td>106 (1.13)</td>
</tr>
<tr>
<td>$B_3$</td>
<td></td>
<td>-0.0385 (0.42)</td>
<td>0.047 (0.50)</td>
<td>-1.168 (1.22)</td>
<td>-0.977 (0.97)</td>
<td>0.039 (0.09)</td>
</tr>
<tr>
<td>$B_4$</td>
<td></td>
<td>-2.002 (-19)</td>
<td>-19 (-2.0)</td>
<td>-0.391 (0.41)</td>
<td>-0.018 (-0.18)</td>
<td>-0.204 (-2.17)</td>
</tr>
<tr>
<td>S.E.</td>
<td></td>
<td>0.01286</td>
<td>0.01285</td>
<td>0.00846</td>
<td>0.00841</td>
<td>504.27</td>
</tr>
<tr>
<td>$a_2$</td>
<td></td>
<td>0 (*)</td>
<td>0 *</td>
<td>-</td>
<td>-</td>
<td>-</td>
</tr>
</tbody>
</table>
| $a_3$  |       | - | - | - | - | 0.62 (+) | - | 0.74 (+) | - | 2.75 (+) | (0.70 (+) | -

(*) Based on (26) or equivalent for $y_t$ and $c_t$.

(+) $H_0$ accepted at 1%.
<table>
<thead>
<tr>
<th>Table 2 - Likelihood Ratio Tests for Model Selection for Actuarial Science (t-statistics in parentheses)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
</tr>
<tr>
<td>Intercepts</td>
</tr>
<tr>
<td>Intercept</td>
</tr>
<tr>
<td></td>
</tr>
<tr>
<td></td>
</tr>
<tr>
<td></td>
</tr>
<tr>
<td></td>
</tr>
<tr>
<td></td>
</tr>
<tr>
<td></td>
</tr>
<tr>
<td></td>
</tr>
</tbody>
</table>

**Working Papers in Economics**

92. M.C. Blad & N. Ulton: Union-Firm Bargaining as a Repeated Prisoner's Dilemma; January 1987
96. B.W. Ross: The Leisure Factor in Entrepreneurial Success during the 'Robber Baron' Era; July 1987
97. F. Gill: Determination of Wage Relativities under the Federal Tribunal: 1953-1974; August 1987
100. L. Haddad: List Revisited: Dynamic Consideration of Trade and Protection; November 1987
101. J. Piggott: General Equilibrium Computation Applied to Public Sector Issues; December 1987
102. J. Carlson & D.W. Findlay: Relative Prices, Wage Indexation and Unemployment; December 1987
104. P. Groenhoven: Taxation and Decentralisation: A Reconsideration of the Costs and Benefits of a Decentralised Tax System; March 1988
106. G. Hills: Satisfactory-Differentiated Tracing Markets: Equilibria under Price Regulation without Entry Restrictions; April 1988
107. B.W. Ross: Strategic Commitment, Unknowledge and the Nature of Entrepreneurial Activity; April 1988
109. B.W. Ross: The Conglomerate and the Focused Agglomeration; Modern Forms of the Leader-Commanded Firm; July 1988
110. L. Ermini: Inertial Behavior on Schedule and Hierarchical Decomposition; July 1988
112. P. Groenhoven: Neo-Classical Value and Distribution Theory; The English Speaking Pioneers; September 1988
<table>
<thead>
<tr>
<th>Working Papers in Economics Published Elsewhere</th>
</tr>
</thead>
<tbody>
<tr>
<td>1. I.G. Sharpe &amp; R.G. Walker</td>
</tr>
<tr>
<td>2. Journal of Accounting Research, 13(2), Autumn 1985</td>
</tr>
<tr>
<td>3. N.V. Lam</td>
</tr>
<tr>
<td>5. V.B. Hall &amp; M.L. King</td>
</tr>
<tr>
<td>6. New Zealand Economic Papers, 10, 1976</td>
</tr>
<tr>
<td>7. A.J. Philips</td>
</tr>
<tr>
<td>8. Economic Record, 53(143), September 1977</td>
</tr>
<tr>
<td>9. N.V. Lam</td>
</tr>
<tr>
<td>11. T.G. Sharpe</td>
</tr>
<tr>
<td>13. W.P. Hogan</td>
</tr>
<tr>
<td>14. Economic Papers, 55, The Economic Society of</td>
</tr>
<tr>
<td>15. Australia and New Zealand</td>
</tr>
<tr>
<td>16. Economic Letters, 2, 1979</td>
</tr>
<tr>
<td>17. Kredit und Kapital, 12(1), 1979</td>
</tr>
<tr>
<td>18. Some Calculations in Stability and Inflation,</td>
</tr>
<tr>
<td>A.H. Bergström et al., (eds.) Wiley &amp; Sons, 1978</td>
</tr>
<tr>
<td>19. Australian Economic Papers, 29(35), December 1980</td>
</tr>
<tr>
<td>20. F.Gill</td>
</tr>
<tr>
<td>22. I.G. Sharpe</td>
</tr>
<tr>
<td>24. R.L. Brown</td>
</tr>
<tr>
<td>25. The Australian Monetary System in the 1970s,</td>
</tr>
<tr>
<td>M. Porter (ed.), Supplement to Economic Board 1978</td>
</tr>
<tr>
<td>26. Australian Economic Record, 56(152), March 1980</td>
</tr>
<tr>
<td>27. Australian Journal of Management, October 1979</td>
</tr>
<tr>
<td>29. W.P. Hogan</td>
</tr>
<tr>
<td>30. Australian Economic Papers, 29(34), June 1980</td>
</tr>
<tr>
<td>32. Australian Economic Papers, 18(33), December 1979</td>
</tr>
<tr>
<td>33. Keynesian Theory, Planning Models, and</td>
</tr>
<tr>
<td>Quantitative Economics, G. Gandolfo</td>
</tr>
<tr>
<td>34. and F. Marzano (eds.), 2, 703-767, 1987</td>
</tr>
<tr>
<td>35. Australian Economic Papers, 21(39), December 1982</td>
</tr>
<tr>
<td>36. Journal of the Operational Research Society,</td>
</tr>
<tr>
<td>(33), 1982</td>
</tr>
<tr>
<td>41. Journal of Industrial Economics, 4(2), March 1983</td>
</tr>
<tr>
<td>42. Review of Economic Studies, 50(160), January 1983</td>
</tr>
<tr>
<td>43. Economic Record, 37(139), December 1981</td>
</tr>
<tr>
<td>44. AFIS, Commission Studies and Selected Papers,</td>
</tr>
<tr>
<td>45. ACS, IV 1982</td>
</tr>
<tr>
<td>46. Economic Record, 50(161), June 1982</td>
</tr>
<tr>
<td>47. Seventh Australian Transport Research</td>
</tr>
<tr>
<td>48. Forum Papers, Hobart, 1982</td>
</tr>
</tbody>
</table>

Copies are available upon request from:
Department of Economics,
The University of Sydney,
N.S.W. 2006, Australia.
56 V.B. hall & P. Saunders
Economic Record, 60(168), March 1984
57 P. Saunders
Economic Record, 59(166), September 1983
58 P Gill
Economic Applicance, 37(3-4), 1984
59 G. Mills & W. Coleman
Journal of Transport Economics and Policy,
16(3), September 1982
60 J. Yates
Economic Papers, Special Edition, April 1983
61 S.S. Joson
Australian Economic Papers, 24(44), June 1985
62 R.T. Ross
Australian Quarterly, 46(3), Spring 1984
63 W.J. Merrilees
Economic Record, 59(165), September 1983
65 A.J. Philp
Australian Economic Papers, 22(41), December 1983
67 V.B. Hall
Economics Letters, 12, 1985
69 V.B. Hall
Energy Economics, 8(2), April 1986
70 F. Gill
Australian Quarterly, 39(2), Winter 1987
71 W.J. Merrilees
Australian Economic Papers, 23(43), December 1984
74 J. Yates
Australian Quarterly, 50(2), Winter 1984
76 V.B. Hall
Economics Letters, 20, 1990
78 S.S. Joson
Journal of Policy Modeling, 8(2), Summer 1986
79 R.T. Ross
Economic Record, 62(178), September 1986
81 R.T. Ross
Australian Bulletin of Labour, 21(4), Sept.1985
82 P. D. Groenewegen
History of Political Economy, 20(4), Winter 1988
84 E.H.A. Gross,
Australian Economic Papers, 27(10), June 1988
94 W.P. Hogan & T.C. Sharpe
Company and Securities Law Journal,
4(1), February 1988
96 B.W. Ross
The Economic and Social Review, 20(1), April 1989
97 F. Gill
Australia's Greatest Asset: Human Resources in
the Nineteenth and Twentieth Centuries,
99 R.T. Ross
Australian Bulletin of Labour,
21(1), December 1988
100 L. Badar
Heta Bulletin, (11), Winter 1989
101 J. Piggott
Public Sector Economics - A Reader,
P.Hare (ed.), Basil Blackwell, 1988
107 D.W. Ross
Prometheus, 6(2), December 1988
108 S.S. Joson
Rivista di Diritto Valutario e DI Economia
Internazionale, 35(2), June 1988
114 V.B. Hall,
T. P. Tran & V.A. Nguyen
116 G. Kingston
Pacific and Asian Journal of Energy, 2(2),
December 1988
122 W.P. Hogan & T.C. Sharpe
Economic Analysis and Policy, 19(1),
March 1989
123 G. Mills
Journal of Transport Economics and Policy,
23, May 1989