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AN EMPIRICAL INVESTIGATION OF SHOCK PERSISTENCE IN ECONOMIC TIME SERIES

Geetha Mayadunne, Merran Evans and Brett Inder

Working Paper No. 18/93

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AN EMPIRICAL INVESTIGATION OF SHOCK PERSISTENCE IN ECONOMIC TIME SERIES

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Whether or not shocks persist has important implications in economics. An empirical study investigates this issue for key Australian and US macroeconomic time series. The existence of persistence is investigated by unit root tests and its magnitude estimated by recently proposed techniques. Results from these different approaches are compared.

1. INTRODUCTION

Despite considerable recent efforts in investigating whether or not shocks to economic time series are persistent, the issue remains inconclusive. This relates to whether time series are better characterized as processes which are stationary around a deterministic trend, or as processes which become stationary on integer or fractional differencing. If the order of differencing (d) required or, equivalently, the order of integration (I(d)) of the time series process is less than one then a shock to the level of the series will eventually die out. If the order of integration is greater than or equal to one, a major implication is that shocks will persist over time.

The persistence of shocks is an important question in economics. In financial market data, the existence of persistence has implications for market efficiency. In macroeconomic data, it has relevance to the rational expectations hypothesis and to the value of counter-cyclical policies. It is important for evaluating the existence of equilibrium relationships that are suggested by economic theory.

Motivated by this debate and its consequences, we investigated persistence in economic time series in terms of both existence and magnitude. Most studies examine the existence of persistence, by way of unit roots tests, but pay little attention to magnitude. There is little value in knowing that shocks persist if the magnitude of persistence is insubstantial. Our objective was to determine whether results from the different approaches are consistent and to identify promising techniques. Our empirical investigation involved a selection of annual and quarterly macroeconomic time series from Australia and the US (from Nelson and Plosser (1982)).

Overall, our investigation shows that it is hard to conclude either on the existence or magnitude of persistence in economic time series. The different methods fail to give consistent results for many of the series. Our findings enhance doubts as to whether economic time series are integrated of order one and suggest that many series might be fractionally integrated, with orders of integration both above and below one. None of the investigated methods seem to be satisfactory in estimating the magnitude of persistence, but the time domain long run variance ratio method appears the most promising.

Persistence is defined and the theoretical background outlined in the following section. The methodology and empirical results are presented in section 3 for the existence of persistence and in section 4 for estimating its magnitude. In section 5 these results are compared.

2. CONCEPTS OF PERSISTENCE

Persistence is defined by the long run effect of a shock on the level of a time series variable continuing indefinitely into the future. It is thus the impact of a unit innovation or shock, e_t, to the variable, $\{y_t\}_{t=1}^{T}$, at time t, on its level at an infinite horizon in time,

$$\lim_{k\to\infty} \frac{\partial y_{t+k}}{\partial e_t}, \qquad (2.1)$$

where y_{t+k} is the level of the variable k periods ahead of period t.

To obtain a measure of this persistence, suppose the first difference of y_t is given the infinite moving average (MA) representation

$$\Delta y_t = (1-L)\, y_t = \mu + \sum_{j=0}^{\infty} a_j L^j e_t = \mu + A(L)e_t, \qquad (2.2)$$

where L is the lag or back shift operator and e_t are independently and identically distributed (iid) random shocks with zero mean and variance σ_e^2. Then the time series can be expressed as

$$y_t = t\mu + a_o \sum_{i=-\infty}^{t} e_i + a_1 \sum_{i=-\infty}^{t-1} e_i + a_2 \sum_{i=-\infty}^{t-2} e_i + \ldots.$$

which gives for k periods ahead,

$$y_{t+k} = (t+k)\mu + a_o \sum_{i=-\infty}^{t+k} e_i + a_1 \sum_{i=-\infty}^{t+k-1} e_i + \ldots + a_k \sum_{i=-\infty}^{t} e_i + \ldots$$

Differentiating this with respect to e_t gives

$$\frac{\partial y_{t+k}}{\partial e_t} = a_o + a_1 + \ldots + a_k \quad .$$

Persistance is measured by this at an infinite time horizon, which is called the cumulative impulse response:

$$A(1) = \lim_{k \to \infty} \frac{\partial y_{t+k}}{\partial e_t} = \sum_{j=0}^{\infty} a_j. \tag{2.3}$$

Hence the infinite sum of the moving average coefficients of the differenced series characterizes the long run effect of a shock on the level of the series.

For a stationary process, y_t can be represented as

$$y_t = \mu + \sum_{j=0}^{\infty} a_j^* L^j e_t = \mu + A^*(L) e_t \tag{2.4}$$

with the usual Wold stationarity conditions on A^*. Then expression (2.2) will be such that

$$A(L) = (1-L)A^*(L)$$

and consequently $A(1) = 0$. Hence a unit shock would have no impact in the long run on the level of a stationary series.

An alternative concept defines persistence as the limit of the variance ratio,

$$V^k = (1+k)^{-1} \frac{\text{var}(y_{t+k+1} - y_t)}{\text{var}(y_{t+1} - y_t)} , \qquad (2.5)$$

or the long run variance ratio,

$$V = \lim_{k \to \infty} V^k . \qquad (2.6)$$

Cochrane (1988) rationalized this as a natural persistence measure as follows. For a random walk, the variance of the (k+1) lagged difference is (k+1) times the variance of the once lagged difference, so irrespective of k V is unity. For a stationary series the variance of the (k+1) lagged difference approaches twice the variance of the original series, so this ratio approaches zero for large k. A process which becomes stationary on differencing can be represented by a random walk and a stationary component, so that the permanent effect of the shock is carried by the random walk and the temporary effect by the stationary component. Hence 1/(1+k) times the variance of the (k+1) lagged difference settles down to the variance of the shock to the permanent component.

This non-parametric measure V can be expressed in the time domain as

$$V = 1 + 2 \sum_{j=1}^{\infty} \rho_j,$$

where ρ_j are the autocorrelations of the differenced process (Δy_t), so it captures all the fluctuations in the process. Equivalently in the frequency domain,

$$V = \frac{2\pi f_{\Delta y}(0)}{var(\Delta y)}, \tag{2.7}$$

where $f_{\Delta y}(0)$ is the spectral density of the differenced process at the zero frequency.

Not surprisingly, these alternative concepts are interrelated.

$$A(1)^2 \frac{\sigma_e^2}{var(\Delta y_t)} = V \tag{2.8}$$

$$f_{\Delta y}(0) = \frac{A(1)^2}{2\pi} \sigma_e^2 \tag{2.9}$$

This enables non-parametric estimation of persistence. Other concepts of persistence are variations of these.

An order of integration of a time series which is greater than or equal to one will result in persistence. If the order of integration is less than one, then shocks should eventually die out. This can be demonstrated with the class of fractionally integrated autoregressive moving average processes, denoted ARFIMA(p,d,q), a generalization of the ARIMA class of Box and Jenkins (1976). [See Granger and Joyeux (1980), Hosking (1981), Geweke and Porter-Hudak (1983) and Sowell (1990, 1992a, 1992b)]. ARFIMA models were introduced to allow for cases where shocks do not die out as quickly as the exponential decay implied by processes which are I(0). They are often termed long memory models.

An ARFIMA(p,d,q) representation of $\{y_t\}_{t=1}^{T}$ is given by

$$\Phi(L)(1-L)^d y_t = \Theta(L)e_t \ , \tag{2.10}$$

where d is a real value, $e_t \sim$ iid $(0,\sigma_e^2)$, $\Phi(L) = 1 - \phi_1 L - \phi_2 L^2 - \ldots - \phi_p L^p$, $\Theta(L) = 1 - \theta_1 L - \theta_2 L^2 - \ldots - \theta_q L^q$, with roots of $\Phi(L) = 0$ and $\Theta(L) = 0$ lying outside the unit circle. For any real number $d > -1$, the operator $(1-L)^d$ is defined by means of the binomial expansion

$$(1-L)^d = \sum_{j=0}^{\infty} \frac{\Gamma(j-d)}{\Gamma(-d)\ \Gamma(j+1)} L^j \ ,$$

where $\Gamma(.)$ is the gamma function. The process is both stationary and invertible when $-0.5 < d < 0.5$. For values of $d > 0.5$ the process is defined by taking a suitable number of integer differences of $\{y_t\}$. Suppose $\{y_t\}$ has the representation

$$(1-L)^d y_t = \sum_{j=0}^{\infty} a_j L^j e_t . \tag{2.11}$$

Then it can be shown that the cumulative impulse response

$$\begin{aligned} A(1) &= \sum_{j=0}^{\infty} a_j && \text{if } d = 1 \\ &= 0 && \text{if } d < 1 \\ &= \infty && \text{if } d > 1. \end{aligned} \tag{2.12}$$

Hence, for a difference stationary process $\sum_{j=0}^{\infty} a_j$ is finite, and for a random walk, $A(1) = 1$. However, when the process is fractional noise, $A(1) = 0$ if $d < 1$, but $A(1) = \infty$ if $d > 1$.

Similar results have been determined for the long run variance ratio persistence measure V. For process (2.11) the spectral density at frequency λ, $f_{\Delta y}(\lambda)$, is given by

$$f_{\Delta y}(\lambda) = \frac{\sigma_e^2}{2\pi} \left|1-e^{-i\lambda}\right|^{-2(d-1)} \left|A(e^{-i\lambda})\right|^2. \qquad (2.13)$$

When $d = 1$, $f_{\Delta y}(0) = \frac{\sigma_e^2}{2\pi} |A(1)|^2$, and $\mathrm{var}(\Delta y_t) = \{\sum_{j=0}^{\infty} a_j^2\} \sigma_e^2$. Therefore, substituting for $f_{\Delta y}(\lambda)$ and $\mathrm{var}(\Delta y_t)$, in (2.7), gives

$$V = \{\sum_{j=0}^{\infty} a_j\}^2 \frac{1}{\sum_{j=0}^{\infty} a_j^2} .$$

It follows that when $\{y_t\}$ is a random walk, $V = 1$. From (2.13), when $d < 1$, $f_{\Delta y}(0) = 0$ and hence $V = 0$, and when $d > 1$, $f_{\Delta y}(0) = \infty$ and $V = \infty$.

Hence, for any process integrated of order d, written I(d), a shock will be persistent if $d \geq 1$ and will die out if $d < 1$.

3. INVESTIGATION OF EXISTENCE OF PERSISTENCE

3.1 Methodology

The well known unit root tests were used to explore the existence of persistence and the fractional integration parameter was estimated to determine its magnitude.

The unit root tests employed were the Augmented Dickey-Fuller (ADF) t-type test of (Dickey and Fuller (1984, 1985)), the Phillips-Perron $Z(\tau)$ test (Perron and Phillips (1987)) and Phillips and Perron (1988) and the Kwiatkowski-Phillips-Shmidt-Shin (KPSS) test (Kwiatkowski et al (1992)). The first two test the null hypothesis of a unit root (d = 1 or I(1)) against the alternative of covariance stationarity (d = 0 or I(0)). The third reverses these hypotheses and tests whether the series is I(0) against I(1). As all are known to be biased towards accepting the null hypothesis, it seemed desirable to conduct tests for each null hypothesis.

No "satisfactory" tests for fractional integration have yet been developed, so we estimated the fractional integration parameter d. Several methods are available: a frequency domain regression method by Geweke and Porter-Hudak (GPH) (1983), maximum likelihood estimation (MLE) in the frequency domain by Fox and Taqqu (1986) and MLE in the time domain by Sowell (1992a, 1992b). Despite the noted superiority of the time domain MLE (see Diebold and Nerlove (1990), Sowell (1991,1992 a, b)) it is computationally very difficult, so the frequency domain two-step regression method of GPH was used. This involves first estimating d ignoring the autoregressive and moving average parameters, then filtering the series $\{y_t\}$ by $(1-L)^d$ and estimating the autoregressive and moving average parameters by the usual Box-Jenkins methodology. Details of the procedure can be found in GPH.

Two data sets, both comprising long annual series on macroeconomic data, were considered in this study. The first relates to the U.S. economy, and was first introduced in Nelson and Plosser (1982). This set of data has been used in a number of recent empirical studies on unit roots tests and persistence. The second set of series uses Australian Bureau of Statistics from 1901-1992.

3.2 Empirical results with economic time series

Results of tests for the existence of persistence are shown in Table 1. For the ADF t-type tests the sequential procedure of Campbell and Perron (1991) was used to select the lag truncation parameter l in the ADF-regression. To adequately approximate the dynamics of the process, a starting value of 10 was chosen, then reduced systematically, until the estimated coefficient of the last included lag was found to be significant. Except for unemployment rates, exchange rate, and bond yield, which included a constant only, all were estimated including the time trend. Supporting the choice of l, residual autocorrelations of the chosen ADF-regression residuals for each series were all near zero. The ADF test statistics show that the null hypothesis of I(1) is rejected only for the US and Australian unemployment rates. Large negative values lead to rejection of the null hypothesis at a 5% level of significance for the US unemployment rate and for Australian at a 10% level. Therefore the test concludes that all except these two series are integrated of order one.

Application of the Phillips-Perron $Z(\tau)$ test requires selections of the lag truncation parameter l for the estimating long run variance of the stochastic component. For almost all the differenced series, insignificant autocorrelations beyond the first few lags indicate a small value. However, following Perron and Phillips (1987) we examined values of $l = 1,2,..,10$. Except the US unemployment rate and industrial production index, the null hypothesis of a unit root is not rejected for all lags. Large negative values for the test statistic mean the null hypothesis is rejected for the US unemployment rate at the 5% level of significance, at all lags, but for industrial production index at the 10% level only for lags 1 to 4.

The KPSS test again involves selection of the lag truncation parameter l for the estimator of the long run variance of the stochastic component. We used the value chosen by Inder's (1992) approximate method: the lag length selected corresponds to the last significant autocorrelation of Δw_t, where $\{w_t\}$ are the residuals from the regression of the series on a constant and if appropriate also a time trend. This leads to the rejection of the null hypothesis of stationarity for all series except the US unemployment rate.

The fractional integration parameter estimates from the GPH method and their asymptotic standard errors, given in the last column of Table 1, suggest the fractionally integrated nature of the series. The US CPI (1.45) and Australian nominal GDP (1.84), nominal consumption (1.38), wages (1.43), M1 (1.54), M3 (1.76) and bond yield (1.41) have fractional integration parameter estimates well above one. US nominal GNP, nominal wages, employment, money stock, velocity, stock prices, bond yield, quarterly GNP, and the Australian unemployment rate, real GDP and real consumption have fractional integration parameter estimates close to one. The estimates for the remaining series are well below one.

Fractional integration parameter estimates greater than 1.5 observed for Australian nominal GDP, M1, and M3 suggest that these series would not reach stationarity on differencing once. The autocorrelations of their first differences take a slightly longer time to decay to zero compared with the remaining series, and hence suggest a possibility that they might be integrated of an order greater than 1.5.

4. ESTIMATION OF MAGNITUDE OF PERSISTENCE

4.1 Methodology

Estimation of persistence based on both the cumulative impulse response function A(1) in (2.13) and the variance ratio V (in (2.5) and (2.6)) involves practical difficulties.

The estimation of A(1) is problematic as it is an infinite sum of moving average coefficients. We followed Campbell and Mankiw (1987a,b) in modelling y_t by an ARIMA(p,1,q) process, or (2.10) with d = 1,

$$\Phi(L)\ \Delta y_t = \Theta(L)e_t + \mu, \qquad (4.1)$$

using Box-Jenkins (1976) methodology. We approximated A(1) by

$$\hat{\Phi}(L)^{-1}\hat{\Theta}(L), \qquad (4.2)$$

where $\Phi(L)$ and $\Theta(L)$ satisfy the conditions of stationarity and invertibility. For a stationary series, differencing would induce a unit root in the moving average component, so we allowed for at least one MA parameter in (4.1) and estimated $\Phi(L)$ and $\Theta(L)$ using exact maximum likelihood.

The time domain variance ratio estimator of (2.5) with large k has the advantage of accounting for autocorrelations between observations far apart. Further, for difference stationary time series $\hat{V}^k$ is the Bartlett estimator of the spectral density at frequency zero, having the limit distribution $\hat{V}^k \sim N(V,\ (V)^2 4k/\ 3T)$. However the choice of k is a problem: two few autocorrelations may obscure trend reversion manifested in higher autocorrelations, but too many may tend to find excessive trend reversion since as k approaches the sample size the estimate trivially approaches zero. For this reason, we estimated V^k with a range of values of k.[1]

Cogley proposed the frequency domain estimator $\hat{V}^f$ for estimating V by (2.7). $\hat{f}_{\Delta y}(0)$ is obtained by smoothing the first few periodogram ordinates of Δy_t so that

$$\hat{f}_{\Delta y}(0) = \sum_{j=1}^{m} W(\omega_j)\, I(\omega_j), \tag{4.3}$$

where $\omega_j = 2\pi j/T$, $j = 1,\ldots,T-1$ and $I(\omega_j)$ denotes the periodogram ordinates estimated by

$$I(\omega_j) = (2\pi T)^{-1} \left| \sum_{t=1}^{T} (\Delta y_t - \hat{\mu}) \exp(-i\omega_j t) \right|. \tag{4.4}$$

$\hat{\mu}$ is the mean of Δy_t. $W(\omega_j)$ is a weight function concentrated near frequency zero, truncated at $\omega = 2\pi m/T$, determined from

$$W_n(\omega_j) = (2\pi k)^{-1} \left\{ \frac{\sin(k\omega_j/2)}{\sin(\omega_j/2)} \right\}^2, \quad j = 1, \ldots m,$$

and normalized so that it sums to one. Cogley proposed setting m to 10. This approach is superior to the time domain method in that a value of k need not be selected to represent the long run, but it is subject to problems inherent in the estimation of spectral densities.

From the long run variance ratio estimates, $\hat{V}^k$ and $\hat{V}^f$, the corresponding persistence estimates $\hat{A}(1)$ were then computed using transformation (2.8). Noting that $\{1-\sigma_e^2/\mathrm{var}(\Delta y)\}$ is the fraction of the variance that is predictable from the past history of the process, Campbell and Mankiw (1987a) proposed estimating it by $\hat{\rho}_1^2$. Thus $\hat{\rho}_1^2$ provides an approximation when the process is AR(1), and otherwise underestimates it.

4.2 Empirical results with economic time series

Persistence estimates from ARMA models were obtained by estimating a range of ARMA(p,q) models for the first difference of the series. Persistence was estimated for each of the combinations of p and q values being 0,1,2. (For Australian nominal GDP and wages q = 3 was also included, as was q = 4 for nominal consumption and M3.) Estimated models which violated conditions of stationarity and/or invertibility and white noise residuals were eliminated. The estimates, reported in Table 2, vary across the models, with a typical moderate variation of the order of about 0.01. Averages were taken for the estimates arising from the different models (excluding any outliers).

All the estimates were well above zero, implying that a shock to any of the series may be permanent. The estimate for US real GNP shows that a 1% shock is likely to increase its level by about 1.46% in the long run, which is in agreement with previous findings. For series with a higher order of integration we would expect shocks to be more persistent. We find higher persistence estimates for US employment, CPI, bond yield, money stock and Australian nominal GDP, nominal consumption, wages, M1, M3. These results are consistent with their higher orders of integration.

There are however several inconsistencies to be noted here. First, the US unemployment rate shows a persistence estimate of about 1.22, whereas the unit root tests and the fractional integration parameter indicate that a shock should die out in the long run suggesting the persistence estimate should be close to zero. Some series that were found to be I(1) by the unit root tests have lower persistence than the US unemployment rate, which suggests that this method may give unreliable estimates of the true magnitude of persistence.

Persistence estimates based on the time and frequency domain long run variance ratio estimates are reported in Table 3. For values of k = 5, 10, 15, 20 and 25, estimates $\hat{V}^k$ were determined and hence the parametric estimates $\hat{A}(1)^k$ of A(1). For the US and Australian unemployment rate they show a steady decline, indicating the possibility of reaching zero at large values of k_1. For the remaining series they increase with k, but for most of the series which have a fractional integration parameter estimate less than one, this increase seems to be marginal, but for those with estimates near or larger than it is more marked. The disturbing feature here again is that series with fractional integration parameter estimates less than one do not show the expected steady decline of $\hat{V}^k$ and $\hat{A}(1)^k$ to zero, and standard errors also are very large.

Frequency domain estimates of persistence $\hat{A}(1)^f$ using $\hat{V}^f$ were computed and shown in the last column of Table 3. There is a weak indication that series with fractional parameter estimates greater than and near one give large values for $\hat{V}^f$ and vice versa. However, unexpectedly, US industrial production and Australian real consumption and exchange rates show smaller values of $\hat{V}^f$ than that of the US unemployment rate. Theoretically estimates of persistence using $\hat{V}^f$ are expected be satisfactory because the long run corresponds to the zero frequency and, at least in theory, the spectral density can be estimated at this zero frequency. However the results do not provide evidence to judge the reliability of the method.

5. CONCLUDING REMARKS

5.1 Comparison of empirical findings

The empirical evidence on the existence of persistence from unit root tests and fractional integration parameter estimates are somewhat ambiguous. The ADF, Phillips-Perron and KPSS tests indicate an order of integration of one for many time series. US CPI, real wages, velocity, stock prices and quarterly GNP and Australian real GDP, nominal GDP, real consumption, nominal consumption, employment, wages, bond yield and M3 fall into this group. The US real GNP, nominal GNP, real per capita GNP, employment, GNP deflator, nominal wages, money stock and bond yield and Australian M1 and exchange rate fall into this group with a lower level of confidence. For the remaining series it is hard to be conclusive. Thus the unit root test results suggest that, except for US and probably Australian unemployment rate and US industrial production index, the series are integrated of order one and hence shocks to these series should persist and converge to non zero finite values. However, the fractional integration parameter estimates indicate otherwise. Many series have fractional parameter estimates well above one: for example, US CPI and Australian nominal GDP, nominal consumption, wages, M1, M3 and bond yield. Some, like US real wages and Australian employment and exchange rate are below one.

Estimates of magnitude of persistence suggest that shocks are permanent in almost all the series, except for US and probably Australian unemployment rates. Persistence estimates obtained from the different ARMA models are sensitive to the parameterizations, but the variation is moderate. The average estimates generally are all well above zero and often are inconsistent with the observed order of integration. For example, the persistence estimate for the US unemployment rate is 1.22, whereas series that were found to be integrated of an order one or greater give persistence estimates lower than this.

Persistence estimates based on non-parametric variance ratio estimates in the time domain show some promising behavior: except the US and Australian unemployment rate, they typically diverge with increasing values of k, and for the unemployment rate converge to a value close to zero. Overall, this method appears to be the most reliable.

The findings on the existence and magnitude of persistence generally are inconsistent. A plausible explanation is that of an aggregation effect: see Korosi, Lovrics and Matyas (1993) for research on this issue. The low power of the unit root tests and the problems associated with fractional parameter estimates make the findings on existence of persistence dubious. However the high estimates of the magnitude of persistence are also suspect, possibly a result of the problems associated with the estimation methods used. In the parametric and in the time domain variance ratio method we attempted to estimate persistence at a time horizon which in no way could approximate infinity. Given a finite amount of data it is hard to increase the time horizon to desirable lengths. In the frequency domain variance ratio method, problems could have arisen from the spectral density estimate, which is hard to estimate accurately especially near the zero frequency.

5.2 Conclusions

Given the results of our empirical analysis it is hard to arrive at a general conclusion on the magnitude of persistence. One clear indication is that the US unemployment rate is very likely to be integrated of an order less than 0.5, and hence is covariance stationary, and a shock effect will eventually die out. For almost all other series, persistence measures vary but very few are near zero. It seems that virtually all series have substantial persistence, with several showing that shocks of 1% have a long run impact in excess of 1%. The implications of this result for economic analysis and econometric methodology are quite significant.

Estimation of the magnitude of persistence using the time domain long run variance ratio estimator appears to be the best technique of those examined. It enables persistence to be tracked at various time horizons, thereby examining the pattern of persistence behavior, and hence allowing conclusions to be drawn about the order of integration of the process. We have not tested the magnitude of persistence however. Definitive conclusions require a better finite sample approximation to its distribution and standard errors obtained under a normal approximation appear very large.

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FOOTNOTES

1. A further problem is observed through simulation experiments that the estimator distribution is highly skewed, so the normal distribution is not a good approximation. In particular the lower tail critical values are too small (Cogley (1991)).

Table 1a

EXISTENCE OF PERSISTENCE

US time series

Lag	ADF test statistic (chosen lag)	Phillips-Perron test statistic 1	2	3	4	KPSS test statistic 0	1	2	3	4	GPH d estimate (st. error)
Real GNP	-2.99 (1)	-2.29	-2.41	-2.35	-2.26	.63	.33	.24	.20	.17	.69 (.35)
Nominal GNP	-2.32 (1)	-1.64	-1.77	-1.81	-1.78	.76	.39	.27	.22	.18	.89 (.35)
RPC GNP	-3.05 (1)	-2.38	-2.49	-2.49	-2.44	.53	.28	.20	.17	.15	.72 (.35)
Industrial production	-3.08 (0)	-3.19	-3.19	-3.21	-3.19	.82	.44	.32	.26	.22	.65 (.29)
Employment	-3.13 (1)	-2.46	-2.55	-2.57	-2.53	.53	.26	.20	.16	.14	1.14 (.32)
Unemployment	-3.22 (1)	-4.21	-4.56	-5.02	-5.15	.46	.25	.18	.15	.13	.55 (.32)
GNP deflator	-2.52 (1)	-2.05	-2.19	-2.27	-2.31	.49	.26	.18	.14	.12	.79 (.32)
CPI	-1.44 (2)	-1.03	-1.18	-1.25	-1.28	1.84	.94	.64	.49	.40	1.45 (.29)
Nominal wage	-2.62 (1)	-1.77	-1.90	-1.96	-1.98	.61	.32	.21	.17	.14	.95 (.35)
Real Wage	-3.05 (1)	-2.47	-2.51	-2.52	-2.49	.95	.51	.36	.29	.25	.71 (.38)
Money stock	-3.08 (1)	-1.82	-2.04	-2.17	-2.25	.44	.23	.16	.12	.10	1.01 (.32)
Velocity	-1.66 (0)	-1.76	-1.78	-1.71	-1.61	1.77	.93	.65	.50	.42	.89 (.29)
Bond yield	-1.82 (0)	1.73	1.58	1.48	1.42	.78	.42	.30	.23	.20	1.16 (.35)
Stock prices	-2.65 (1)	-2.16	-2.17	-2.15	-2.08	1.23	.64	.45	.36	.30	.93 (.29)
Quarterly GNP	-2.85 (2)	-1.87	-2.05	-2.13	-2.16	1.87	.96	.65	.50	.41	.94 (.24)

Table 1b

EXISTENCE OF PERSISTENCE

Australian time series

Lag	ADF test statistic (chosen lag)	Phillips-Perron test statistic 1	2	3	4	KPSS test statistic 0	1	2	3	4	GPH d estimate (st. error)
Real GDP	-1.23 (4)	-1.53	-1.54	-1.56	-1.49	1.88	.98	.67	.51	.42	.99 (.32)
Nominal GDP	-0.74 (1)	- .22	- .36	- .44	- .48	2.00	1.02	.69	.53	.43	1.84 (.32)
Real consumption	-1.79 (1)	-2.22	-2.25	-2.30	-2.35	1.55	.84	.58	.45	.38	.89 (.32)
Nominal consumption	- .49 (1)	- .14	- .29	- .40	- .48	1.94	.99	.67	.51	.42	1.38 (.32)
Employment	-2.25 (3)	-2.04	-2.24	-2.29	-2.30	1.18	.62	.43	.34	.29	.66 (.32)
Unemployment	-2.73 (1)	-1.88	-1.97	-2.01	-2.01	.77	.41	.28	.22	.19	1.08 (.32)
Wages	-0.98 (2)	.15	- .09	- .21	- .29	1.52	.78	.54	.41	.34	1.43 (.32)
M1	-1.93 (2)	-1.38	-1.52	-1.63	-1.72	.79	.41	.28	.22	.18	1.54 (.32)
M3	1.13 (0)	.98	- .86	- .74	- .63	1.77	.92	.63	.49	.40	1.76 (.32)
Bond yield	- .20 (2)	- .37	- .38	- .39	- .40	4.59	2.35	1.61	1.23	1.01	1.41 (.32)
Exchange rate	- .65 (2)	- .84	- .90	- .95	- .99	7.94	4.09	2.80	2.15	1.75	.37 (.32)

Table 2a

PERSISTENCE ESTIMATES (A(1)) USING TIME DOMAIN ARMA METHODS

US time series

Series	AR(1)	AR(2)	MA(1)	MA(2)	ARMA(1,1)	ARMA(2,1)	ARMA(1,2)	ARMA(2,2)	average
Real GNP	1.52	1.40	1.32	1.51	1.47	1.52	-	-	1.46
Nominal GNP	1.78	1.57	1.44	1.60	1.66	1.78	1.67	-	1.64
RPC.GNP	1.50	1.38	1.31	1.49	1.45	1.49	-	-	1.44
Industrial production	1.04	.93	1.05	.88	1.05	-	.42	.45	.99
Employment	1.47	1.26	1.39	1.35	1.35	-	-	.34	1.36
Unemployment	1.29	1.08	1.37	1.14	1.25	1.16	.89	-	1.22
GNP deflator	1.77	1.79	1.38	1.57	1.82	-	1.82	-	1.69
CPI	2.39	-	1.66	1.92	2.06	-	-	-	2.01
Nominal wages	1.87	1.62	1.47	1.65	1.71	1.57	1.69	1.57	1.64
Real wages	1.24	1.16	1.20	1.19	1.20	-	1.20	-	1.20
Money stock	2.65	2.33	1.58	1.91	2.34	2.55	2.37	2.39	2.27
Velocity	1.12	1.05	1.12	.90	1.13	1.16	.80	1.08	1.05
Bond yield	1.28	2.21	.87	1.43	-	3.88	-	2.91	2.10
Stock prices	1.29	1.08	1.31	1.13	1.23	.74	.79	.93	1.08
Quarterly GNP	1.59	1.84	1.27	1.58	1.75	1.79	1.78	1.53	1.64

Table 2b

PERSISTENCE ESTIMATES (A(1)) USING TIME DOMAIN ARMA METHODS

Australian time series

Series	AR(1)	AR(2)	MA(1)	MA(2)	ARMA(1,1)	ARMA(2,1)	ARMA(1,2)	ARMA(2,2)	average
Real GDP	.90	.96	.89	.96	1.00	.91	1.04	-	.95
Nominal GDP	2.12	2.56	1.39	1.74	2.70	2.45	2.62	2.61	-
Real consumption	.74	.74	.69	.81	.74	-	.82	-	.76
Nominal consumption	2.09	2.53	1.41	1.61	2.95	3.05	3.00	3.60	-
Employment	1.31	1.59	1.16	1.56	1.41	1.64	1.39	-	1.44
Unemployment	.74	1.42	1.33	1.53	1.53	1.43	1.52	1.36	1.36
Wages	3.68	3.03	1.73	2.26	3.15	3.20	2.05	-	-
M1	1.66	1.84	1.32	1.49	2.15	2.16	2.32	2.14	1.89
M3	1.47	1.80	1.25	1.38	3.38	3.30	3.63	3.23	-
Bond yield	1.09	.90	1.14	.91	.30	0.94	0.97	1.31	.95
Exchange rate	1.11	.86	1.21	0.69	1.14	-	.54	.69	1.00

Table 3a

PERSISTENCE ESTIMATES (A(1)) USING VARIANCE RATIO METHODS

U.S. time series

	$\hat{A}(1)^k$					$\hat{A}(1)^f$
k =	5	10	15	20	25	
	Time domain					Spectral domain
Real GNP	1.54	1.75	1.94	2.19	2.43	.88
Nominal GNP	1.77	2.14	2.36	2.55	2.80	1.16
RPC GNP	1.37	1.37	1.36	1.50	1.62	.87
Industrial production	1.26	1.44	1.64	1.87	2.03	.57
Employment	1.47	1.68	1.80	1.94	2.06	.86
Unemployment	1.11	.97	.82	.70	.64	.74
GNP deflator	1.73	2.05	2.31	2.44	2.54	1.44
CPI	1.76	1.89	1.99	2.02	2.03	2.00
Nominal wage	1.86	2.33	2.66	2.88	3.00	1.13
Real wage	1.45	1.75	2.03	2.33	2.57	.83
Money stock	2.46	3.20	3.69	4.10	4.55	1.14
Velocity	1.03	1.02	1.41	1.18	1.21	.92
Bond yield	1.19	1.19	1.32	1.30	1.16	2.07
Stock prices	1.42	1.09	1.20	1.21	1.20	.88
Quarterly GNP	1.86	2.32	2.69	2.94	3.20	1.08

Table 3b

PERSISTENCE ESTIMATES (A(1)) USING VARIANCE RATIO METHODS

Australian time series

	$\hat{A}(1)^k$					$\hat{A}(1)^f$
k =	5	10	15	20	25	
	Time domain					Spectral domain
Real GNP	.97	1.05	1.20	1.38	1.50	1.87
Nominal GNP	1.68	2.18	2.59	3.03	3.37	1.86
Real consumption	2.12	2.86	3.38	3.72	3.98	.82
Nominal consumption	2.35	3.19	3.78	4.20	4.52	.65
Wages	2.90	3.67	4.17	4.59	4.96	2.25
Employment	1.70	2.03	2.37	2.76	3.04	.83
Unemployment	1.31	1.24	1.14	1.04	.93	1.11
M1	1.99	2.63	3.04	3.27	3.46	1.41
M3	2.04	2.76	3.27	3.64	3.94	1.33
Bond yield	1.02	1.16	1.23	1.22	1.16	1.14
Exchange rate	.92	.78	.80	.76	.80	.56