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THE SOURCES AND NATURE OF LONG-TERM MEMORY IN THE BUSINESS CYCLE

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ABSTRACT

This paper examines the stochastic properties of aggregate macroeconomic time series from the standpoint of fractionally integrated models, focusing on the persistence of economic shocks. We develop a simple macroeconomic model that exhibits long-range dependence, a consequence of aggregation in the presence of real business cycles. We then derive the relation between properties of fractionally integrated macroeconomic time series and those of microeconomic data and discuss how fiscal policy may alter the stochastic behavior of the former. To implement these results empirically, we employ a test for fractionally integrated time series based on the Hurst-Mandelbrot rescaled range. This test, which is robust to short-term dependence, is applied to quarterly and annual real GNP to determine the sources and nature of long-term dependence in the business cycle.

1. Introduction

Questions about the persistence of economic shocks currently occupy an important place in macroeconomics. Most of the controversy has centered on whether aggregate time series are better approximated by fluctuations around a deterministic trend or by a random walk plus a stationary or temporary component. The empirical results from these studies are mixed, perhaps because measuring low-frequency components is difficult. Looking at the class of fractionally integrated processes, which exhibits an interesting type of long-range dependence in an elegant and parsimonious way, can help to resolve the problem. This new approach also accords well with the classic NBER business cycle program developed by Wesley Claire Mitchell, who urged examination of trends and cycles at all frequencies.

Economic life does not proceed smoothly: There are good times and bad times, a rhythmical pattern of prosperity and depression. Recurrent downturns and crises take place roughly every three to five years and thus seem part of a nonperiodic cycle. Studying such cycles in detail has been the main activity of twentieth century macroeconomics. Even so, isolating cycles of these frequencies has been difficult because the data evince many other cycles of longer and shorter duration. Mitchell (1927, p. 463) remarks, "Time series also show that the cyclical fluctuations of most (not all) economic processes occur in combination with fluctuations of several other sorts: secular trends, primary and secondary, seasonal variations, and irregular fluctuations." Properly eliminating these other influences has always been controversial. No less an authority than Irving Fisher (1925) considered the business cycle to be a myth, akin to a run of luck at Monte Carlo. In a similar vein, Slutzky (1937) suggested that cycles arise from smoothing procedures used to create the data.

A similar debate is now taking place. The standard methods of removing a linear or exponential trend assume implicitly that business cycles are fluctuations around a trend. Other work (e.g., Nelson and Plosser [1982]) challenges this assumption and posits stochastic trends similar to random walks, highlighting the distinction between temporary and permanent changes. Since the cyclical, or temporary, component is small relative to the fluctuation in the trend component (the random walk part) when viewed empirically, business cycles look more like Fisher's myth. This is important for forecasting purposes, because permanent changes (as in the case of a random walk) have a large effect many periods later, whereas temporary changes (as in stationary fluctuations around a trend) have small future effects. The large random walk component also provides evidence against some theoretical models of aggregate output. Models that focus on monetary or aggregate demand disturbances as a source of transitory fluctuations cannot explain much output variation; supply-side or other models must be invoked (see Nelson and Plosser [1982] and Campbell and Mankiw [1987]).

The recent studies posit a misleading dichotomy, however. In stressing trends versus random walks, they overlook earlier work by Mitchell (1927), Adelman (1965), and Kuznets (1965), who focused on correlations in the data that fall between secular trends and transitory fluctuations. In the language of the early NBER, most recent studies miss Kondratiev, Kuznets, and Juglar cycles. The longer-run (lower-frequency) properties can be difficult to handle with conventional ARMA or ARIMA models because such properties involve what seem to be an excessive number of free parameters. Of course, an MA(120) fits the post-Civil War annual data quite well, but most of the relations would be spurious, and it is doubtful how well such an overfitted specification could predict. Fractionally differenced processes exhibit long-run dependence by adding only one free parameter, the degree of differencing, and show promise in explaining the lower-frequency effects (i.e., Kuznets' (1965) and Adelman's (1965) "long swings," or the effects that persist from one business cycle to the next). Standard methods of fitting Box-Jenkins models have trouble with the number of free parameters needed for long-term dependence, especially the sort captured by a fractional process. We think a better approach is a more direct investigation of this alternative class of stochastic processes.

This paper examines the stochastic properties of aggregate output from the standpoint of fractionally integrated models. We introduce this type of process in section 2 and review its main properties, its advantages, and its weaknesses. Section 3 develops a simple macroeconomic model that exhibits long-term dependence. Section 4 employs a new test for fractional integration in time series to search for long-term dependence in the data. Though related to a test developed by Hurst and Mandelbrot, our model is robust to short-term dependence. Section 5 summarizes and concludes.

2. Review of Fractional Techniques in Statistics

A random walk can model time series that look cyclic but nonperiodic. The first differences of that series (or in continuous time, the derivative) should then be white noise. This is an example of the common intuition that differencing (differentiating) a time series makes it "rougher," whereas summing (integrating) makes it "smoother." Many macroeconomic time series resemble neither a random walk nor white noise, suggesting that some compromise or hybrid between the random walk and its integral may be useful. Such a concept has been given content through the development of the fractional calculus, i.e., differentiation and integration to non-integer orders.[1] The fractional integral of order between zero and one may be viewed as a filter that smooths white noise to a lesser degree than the ordinary integral; it yields a series that is rougher than a random walk but smoother than white noise. Granger and Joyeux (1980) and Hosking (1981) develop the time-series implications of fractional differencing in discrete time. For expositional purposes, we review the more relevant properties in sections 2.1 and 2.2.

2.1. Fractional Differencing

Perhaps the most intuitive exposition of fractionally differenced time series is via their infinite-order autoregressive (AR) and moving-average (MA) representations. Let X_t satisfy

$$(1-L)^d X_t = \epsilon_t, \qquad (2.1)$$

where ϵ_t is white noise, d is the degree of differencing, and L denotes the lag operator. If d = 0, then X_t is white noise, whereas if d = 1, X_t is a random walk. However, as Granger and Joyeux (1980) and Hosking (1981) have shown, d need not be an integer. From the binomial theorem, we have the relation

$$(1-L)^d = \sum_{k=0}^{\infty} (-1)^k \binom{d}{k} L^k, \tag{2.2}$$

where the binomial coefficient $\binom{d}{k}$ is defined as

$$\binom{d}{k} \equiv \frac{d(d-1)(d-2)\cdots(d-k+1)}{k!} \tag{2.3}$$

for any real number d and non-negative integer k.[2] From (2.2), the AR representation of X_t is apparent:

$$A(L)X_t = \sum_{k=0}^{\infty} A_k L^k X_t = \sum_{k=0}^{\infty} A_k X_{t-k} = \epsilon_t, \tag{2.4}$$

where $A_k \equiv (-1)^k \binom{d}{k}$. The AR coefficients are often reexpressed more directly in terms of the gamma function:

$$A_k \equiv (-1)^k \binom{d}{k} = \frac{\Gamma(k-d)}{\Gamma(-d)\Gamma(k+1)} . \tag{2.5}$$

By manipulating (2.1) mechanically, X_t may also be viewed as an infinite-order MA process, since

$$X_t = (1-L)^{-d}\epsilon_t = B(L)\epsilon_t \text{ and } B_k = \frac{\Gamma(k+d)}{\Gamma(d)\Gamma(k+1)} . \tag{2.6}$$

The particular time-series properties of X_t depend intimately on the value of the differencing parameter d. For example, Granger and Joyeux (1980) and Hosking (1981) show that X_t is stationary when d is less than one-half, and invertible when d is greater than minus one-half. Although the specification in (2.1) is a fractional integral of pure white noise, the extension to fractional ARIMA models is clear.[3]

The AR and MA representations of fractionally differenced time series have many applications and illustrate the central properties of fractional processes, particularly long-term dependence. The MA coefficients B_k give the effect of a shock k periods ahead and indicate the extent to which current levels of the process depend on past values. How fast this dependence decays furnishes valuable information about the process. Using Stirling's approximation, we have

$$B_k \approx \frac{k^{d-1}}{\Gamma(d)} \tag{2.7}$$

for large k. Comparing this with the decay of an AR(1) process highlights a central feature of fractional processes: They decay hyperbolically, at rate k^{d-1}, rather than at the exponential rate of ρ^k for an AR(1). For example, compare in figure 1 the autocorrelation function of the fractionally differenced series $(1-L)^{0.475}X_t = \epsilon_t$ with that of the AR(1) X_t $0.9X_t + \epsilon_t$. Although they both have first-order autocorrelations of 0.90, the AR(1)'s autocorrelation function decays much more rapidly.

Figure 2 plots the impulse-response functions of these two processes. At lag 1, the MA coefficients of the fractionally differenced series and the AR(1) are 0.475 and 0.900, respectively. At lag 10, these coefficients are 0.158 and 0.349, while at lag 100, they fall to 0.048 and 0.000027. The persistence of the fractionally differenced series is apparent at the longer lags. Alternatively, we may ask what value of an AR(1)'s autoregressive parameter will yield, for a given lag, the same impulse response as the fractionally differenced series (2.1). This value, simply the k-th root of B_k, is plotted in figure 3 for various lags when $d = 0.475$. For large k, this autoregressive parameter must be very close to unity.

These representations also show how standard econometric methods can fail to detect fractional processes, necessitating the methods described in section 4. Although a high-order ARMA process can mimic the hyperbolic decay of a fractionally differenced series in finite samples, the large number of parameters required would give the estimation a poor rating from the usual Akaike or Schwartz criteria. An explicitly fractional process, however, captures that pattern with a single parameter, d. Granger and Joyeux (1980) and Geweke and Porter-Hudak (1983) provide empirical support for this by showing that fractional models often outpredict fitted ARMA models.

The lag polynomials A(L) and B(L) provide a metric for the persistence of X_t. Suppose X_t represents GNP, which falls unexpectedly this year. How much should this alter a forecast of GNP? To address this issue, define C_k as the coefficients of the lag polynomial C(L) that satisfy the relation $(1-L)X_t = C(L)\epsilon_t$, where the process X_t is given by (2.1). One measure used by Campbell and Mankiw (1987) is

$$\lim_{k\to\infty} B_k = \sum_{k=0}^{\infty} C_k = C(1). \tag{2.8}$$

For large k, the value of B_k measures the response of X_{t+k} to an innovation at time t, a natural metric for persistence.[4] From (2.7), it is immediate that for $0 < d < 1$, $C(1) = 0$, and that asymptotically, there is no persistence in a fractionally differenced series, even though the autocorrelations die out very slowly.[5] This holds not only for $d < 1/2$ (the stationary case), but also for $1/2 < d < 1$ (the nonstationary case).

From these calculations, it is apparent that the long-run dependence of fractional processes relates to the slow decay of the autocorrelations, not to any permanent effect. This distinction is important; an IMA(1,1) can have small yet positive persistence, but the coefficients will never mimic the slow decay of a fractional process.

The long-term dependence of fractionally differenced time series forces us to modify some conclusions about decomposing time series into "permanent" and "temporary" components. Although Beveridge and Nelson (1981) show that nonstationary time series can always be expressed as the sum of a random walk and a stationary process, the stationary component may exhibit long-range dependence. This suggests that the temporary component of the business cycle may be transitory only in the mathematical sense and that it is, for all practical purposes, closer to what we think of as a long, nonperiodic cycle.

2.2. Spectral Representation

The spectrum, or spectral density (denoted $f(\omega)$), of a time series specifies the contribution each frequency makes to the total variance. Granger (1966) and Adelman (1965) have pointed out that most aggregate economic time series have a typical spectral shape, where the spectrum increases dramatically as the frequency approaches zero ($f(\omega) \to \infty$ as $\omega \to 0$). Most of the power (variance) seems to be concentrated at low frequencies. However, prewhitening or differencing the data often leads to overdifferencing, or "zapping out" the low-frequency component, and frequently replaces the peak by a dip at zero. Fractional differencing yields an intermediate result. The spectra of fractional processes exhibit peaks at zero (unlike the flat spectrum of an ARMA process), but ones not so sharp as those of a random walk. A fractional series has a spectrum that is richer in low-frequency terms and that shows more persistence. We illustrate this by calculating the spectrum of fractionally integrated white noise, and present several formulas needed in sections 3 and 4. Given $X_t = (1-L)^{-d}\epsilon_t$, the series is clearly the output of a linear system with a white noise input, so that the spectrum of X_t is[6]

$$f(\omega) = \frac{1}{|1-z|^{2d}} \frac{\sigma^2}{2\pi},$$

where $z \equiv e^{i\omega}$, and $\sigma^2 \equiv E[\epsilon_t^2]$. (2.9)

The identity $|1-z|^2 = 2(1-\cos(\omega))$ implies that for small ω,

$$f(\omega) = c\omega^{-2d}, \quad c \equiv \frac{\sigma^2}{2\pi}. \tag{2.10}$$

This approximation encompasses the two extremes of white noise (or a finite ARMA process) and a random walk. For white noise, $d = 0$ and $f(\omega) = c$, while for a random walk, $d = 1$ and the spectrum is inversely proportional to ω^2. A class of processes of current interest in the statistical physics literature, called 1/f noise, matches fractionally integrated noise with $d = 1/2$.

3. A Simple Macroeconomic Model with Long-Term Dependence

Over half a century ago, Wesley Claire Mitchell (1927, p. 230) wrote that "We stand to learn more about economic oscillations at large and about business cycles in particular, if we approach the problem of trends as theorists, than if we confine ourselves to strictly empirical work." Indeed, gaining insights beyond stylized facts requires guidance from theory. Theories of long-range dependence may provide organization and discipline in constructing models of growth and business cycles. They can also guide future research by predicting policy effects, postulating underlying causes, and suggesting new ways to analyze and combine data. Ultimately, examining the facts serves only as a prelude. Economic understanding requires more than a consensus on the Wold representation of GNP; it demands a falsifiable model based on the tastes and technology of the actual economy.

Thus, before testing for long-run dependence, we develop a simple model in which aggregate output exhibits long-run dependence. The model presents one reason that macroeconomic data might show the particular stochastic structure for which we test. It also shows that models can restrict the fractional differencing properties of time series, thus holding promise for distinguishing between competing theories. Furthermore, the maximizing model presented below connects long-term dependence to the central economic concepts of productivity, aggregation, and the limits of the representative-agent paradigm.

3.1. A Simple Real Model

One plausible mechanism for generating long-run dependence in output, which we will mention briefly and not pursue, is that production shocks themselves follow a fractionally integrated process. This explanation for persistence follows that used by Kydland and Prescott (1982). In general, such an approach begs the question, but in the present case, evidence from geophysical and meteorological records suggests that many economically important shocks have long-run correlation properties. Mandelbrot and Wallis (1969b), for instance, find long-run dependence in rainfall, river flows, earthquakes, and weather patterns (as measured by tree rings and sediment deposits).

A more satisfactory model explains the time-series properties of data by producing them despite white noise shocks. This section develops such a model with long-run dependence, using a linear quadratic version of the real business cycle model of Long and Plosser (1983) and the aggregation results of Granger (1980). In our multisector model, the output of each industry (or island) follows an AR(1) process, but aggregate output with N sections follows an ARMA (N,N-1) process, making dynamics with even a moderate number of sectors unmanageable. Under fairly general conditions, however, a simple fractional process can closely approximate the true ARMA specification.

Consider a model economy with many goods and a representative agent who chooses a production and consumption plan. The infinitely lived agent inhabits a linear quadratic version of the real business cycle model and has a lifetime utility function of $U = \sum\beta^t u(C_t)$, where C_t is an Nx1 vector denoting period t consumption of each of the N goods in our economy. Each period's utility function $u(C_t)$ is given by

$$u(C_t) = C_t'\iota - \frac{1}{2}C_t'BC_t, \tag{3.1}$$

where ι is an Nx1 vector of ones. In anticipation of the aggregation considered later, we assume B to be diagonal so that $C_t'BC_t = \sum b_{ii}C_{it}^2$. The agent faces a resource constraint: Total output Y_t may be either consumed or saved. Thus,

$$C_t + S_t\iota = Y_t, \tag{3.2}$$

where the i,j-th entry S_{ijt} of the NxN matrix S_t denotes the quantity of good j invested in process i at time t, and where it is assumed that any good Y_{jt} may be consumed or invested. Output is determined by the random linear technology

$$Y_t = AS_t + \epsilon_t, \tag{3.3}$$

where ϵ_t is a (vector) random production shock whose value is realized at the beginning of period t+1. The matrix A consists of the input-output parameters a_{ij}. To focus on long-term dependence, we restrict A's form. Thus, each sector uses only its own output as input, yielding a diagonal A matrix and allowing us to simplify the notation by defining $a_i \equiv a_{ii}$. This diagonal case might occur, for example, when a number of distinct islands are producing different goods. To further simplify the problem, we assume that all commodities are perishable and that capital depreciates at a rate of 100 percent. Since the state of the economy in each period is fully specified by that period's output and productivity shock, it is useful to denote that vector $Z_t \equiv [Y_t' \ \epsilon_t']'$.

Subject to the production function (3.3) and the resource constraint (3.2), the agent maximizes expected lifetime utility as follows:

$$\underset{\{S_t\}}{\mathrm{Max}}\ E[U|Z_t] = \underset{\{S_t\}}{\mathrm{Max}} E\left[\sum_{\tau=t}^{\infty} \beta^{\tau-t} u(Y_t - S_t \iota) \,\middle|\, Z_t\right], \tag{3.4}$$

where we have substituted for consumption in (3.4) using the budget equation (3.2). This maps naturally into a dynamic programming formulation, with a value function $V(Z_t)$ and optimality equation

$$V(Z_t) = \underset{\{S_t\}}{\mathrm{Max}} \left\{ u(Y_t - S_t\iota) + \beta E[V(Z_{t+1})|Z_t] \right\}. \tag{3.5}$$

With quadratic utility and linear production, it is straightforward to discover and verify the form of $V(Z_t)$:

$$V(Y,\epsilon) = q'Y + Y'PY + R + E[\epsilon'T\epsilon], \tag{3.6}$$

where q and R denote Nx1 vectors and P and T are NxN matrices, with entries being fixed constants given by the matrix Riccati equation resulting from the value function's recursive definition.[7] Given the value function, the first-order conditions of the optimality equation (3.5) yield the chosen quantities of consumption and investment/savings and, for the example presented here, have the following closed-form solutions:

$$S_{it} = \frac{b_i}{b_i - 2\beta P_i a_i^2} Y_{it} + \frac{\beta q_i a_i - 1}{b_i - 2\beta P_i a_i^2} \tag{3.7}$$

and

$$C_{it} = \frac{2\beta P_i a_i^2}{2\beta P_i a_i^2 - b_i} + \frac{\beta q_i a_i - 1}{2\beta P_i a_i^2 - b_i}, \tag{3.8}$$

where

$$P_i = b_i \left[\frac{a_i - \sqrt{(1 + 4\beta)a_i^2 - 4}}{4\beta a_i} \right] . \tag{3.9}$$

The simple form of the optimal consumption and investment decision rules comes from the quadratic preferences and the linear production function. Two qualitative features bear emphasizing. First, higher output today will increase both current consumption and current investment, thus increasing future output. Even with 100 percent depreciation, no durable commodities, and i.i.d. production shocks, the time-to-build feature of investment induces serial correlation. Second, the optimal choices do not depend on the uncertainty that is present. This certainty equivalence feature is clearly an artifact of the linear-quadratic combination.

The time series of output can now be calculated from the production function (3.1) and the decision rule (3.7). Quantity dynamics then come from the difference equation

$$Y_{it+1} = \frac{a_i b_i}{b_i - 2\beta P_i a_i^2} Y_{it} + K_i + \epsilon_{it+1}, \tag{3.10}$$

or

$$Y_{it+1} = \alpha_i Y_{it} + K_i + \epsilon_{it+1}, \tag{3.11}$$

where K_i is some fixed constant. The key qualitative property of quantity dynamics summarized by (3.11) is that output Y_{it} follows an AR(1) process.

Higher output today implies higher output in the future. That effect dies off at a rate that depends on the parameter α_i, which in turn depends on the underlying preferences and technology.

The simple output dynamics for a single industry or island neither mimics business cycles nor exhibits long-run dependence. However, aggregate output, the sum across all sectors, does show such dependence, which we demonstrate here by applying the aggregation results of Granger (1980, 1988).

It is well known that the sum of two series X_t and Y_t, each AR(1) with independent error, is an ARMA(2,1) process. Simple induction then implies that the sum of N independent AR(1) processes with distinct parameters has an ARMA(N,N-1) representation. With more than six million registered businesses in America (Council of Economic Advisors, 1988), the dynamics can be incredibly rich -- and the number of parameters unmanageably huge. The common response to this problem is to pretend that many different firms (islands) have the same AR(1) representation for output, which reduces the dimensions of the aggregate ARMA process. This "canceling of roots" requires identical autoregressive parameters. An alternative approach reduces the scope of the problem by showing that the ARMA process approximates a fractionally integrated process and thus summarizes the many ARMA parameters in a parsimonious manner. Though we consider only the case of independent sectors, dependence is easily handled.

Consider the case of N sectors, with the productivity shock for each serially uncorrelated and independent across islands. Furthermore, let the sectors differ according to the productivity coefficient a_i. This implies differences in α_i, the autoregressive parameter for sector i's output Y_{it}. One of our key results is that under some distributional assumptions about α_i's aggregate output, Y_t^a follows a fractionally integrated process, where

$$Y_t^a \equiv \sum_{i=1}^{N} Y_{it}. \tag{3.12}$$

To show this, we approach the problem from the frequency domain and apply spectral methods, which often simplify problems of aggregation.[8] Let $f(\omega)$ denote the spectrum (spectral density function) of a random variable, and let $z = e^{-i\omega}$. From the definition of the spectrum as the Fourier transform of the autocovariance function, the spectrum of Y_{it} is

$$f_i(\omega) = \frac{1}{|1 - \alpha_i z|^2} \frac{\sigma_i^2}{2\pi}. \tag{3.13}$$

Similarly, independence implies that the spectrum of Y_t^a is

$$f(\omega) = \sum_{i=1}^{N} f_i(\omega). \tag{3.14}$$

The α_i's measure an industry's average output for given input. This attribute of the production function can be thought of as a drawing from nature, as can the variance of the productivity shocks ϵ_{it} for each sector. Thus, it makes sense to think of the a_i's as independently drawn from a distribution G(a) and the α_i's as drawn from $F(\alpha)$. Provided that the ϵ_{it} shocks are independent of the distribution of α_i's, the spectral density of the sum can be written as

$$f(\omega) = \frac{N}{2\pi} E[\sigma^2] \cdot \int \frac{1}{|1-\alpha z|^2} dF(\alpha). \tag{3.15}$$

If the distribution $F(\alpha)$ is discrete, so that it takes on m (< N) values, Y_t^1 will be an ARMA (m, m-1) process. A more general distribution leads to a process that no finite ARMA model can represent. To further specify the process, take a particular distribution for F, in this case a variant of the beta distribution.[9] In particular, let α^2 be distributed as beta (p,q), which yields the following density function for α:

$$dF(\alpha) = \begin{cases} \frac{2}{\beta(p,q)} \alpha^{2p-1} (1-\alpha^2)^{q-1} d\alpha & \text{if } 0 \le \alpha \le 1 \\ 0 & \text{otherwise}, \end{cases} \tag{3.16}$$

with $(p,q) > 0$.[10]

Obtaining the Wold representation of the resulting process requires a little more work. First, note that

$$1/|1 - \alpha z|^2 = \frac{1}{2(1 - \alpha^2)} \left[\frac{1 + \alpha z}{1 - \alpha z} + \frac{1 + \alpha \bar{z}}{1 - \alpha \bar{z}} \right], \tag{3.17}$$

where $\bar{z}$ denotes the complex conjugate of z, and the terms in brackets can be further expanded by long division. Substituting this expansion and the beta distribution (3.16) into the expression for the spectrum and simplifying (using the relation $z + \bar{z} = 2\cos(\omega)$) yields

$$f(\omega) = \int_0^1 \left[2 + 2 \sum_{k=1}^{\infty} \alpha^k \cos(k\omega) \right] \frac{2}{\beta(p,\ 1)} \alpha^{2p-1} (1 - \alpha^2)^{q-2} d\alpha. \tag{3.18}$$

Then, the coefficient of $\cos(k\omega)$ is

$$\int_0^1 \frac{2\alpha^k}{\beta(p,\ q)} \alpha^{2p+k-1} (1 - \alpha^2)^{q-2}\, d\alpha. \tag{3.19}$$

Since the spectral density is the Fourier transform of the autocovariance function, (3.19) is the k-th autocovariance of Y_t^a. Furthermore, because the integral defines a beta function, (3.19) simplifies to $\beta(p+k/2, q - 1)/ \beta(p,q)$. Dividing by the variance gives the autocorrelation coefficients, which reduce to

$$\rho(k) = \frac{\Gamma(p + q - 1)}{\Gamma(p)} \frac{\Gamma(p + \frac{k}{2})}{\Gamma(p + \frac{k}{2} + q + 1)}\ . \tag{3.20}$$

Using the result from Stirling's approximation $\Gamma(a+k)/\Gamma(b+k) \approx k^{a-b}$, (3.20) is proportional (for large lags) to k^{1-q}. Thus, aggregate output Y_t^1 follows a fractionally integrated process of the order $d = 1 - \frac{q}{2}$.

Furthermore, as an approximation for long lags, this does not necessarily rule out interesting correlations at higher, e.g., business cycle, frequencies. Similarly, comovements can arise as the fractionally integrated income process induces fractional integration in other observed time series. This phenomenon has been generated by a maximizing model based on tastes and technologies.[11]

In principle, all of the model's parameters may be estimated, from the distribution of production functions to the variance of output shocks. Although to our knowledge no one has explicitly estimated the distribution of production function parameters, many people have estimated production functions across industries.[12] (One of the better recent studies disaggregates to 45 industries.[13]) For our purposes, the quantity closest to a_i is the value-weighted intermediate-product factor share. Using a translog production function, this gives the factor share of inputs coming from industries, excluding labor and capital. These inputs range from a low of 0.07 for radio and television advertising to a high of 0.81 for petroleum and coal products. Thus, even a small amount of disaggregation reveals a large dispersion, suggesting the plausibility and significance of the simple model presented in this section.

Although the original motivation for our real business cycle model was to illustrate how long-range dependence could arise naturally in an economic system, our results have broader implications for general macroeconomic modeling. They show that moving to a multiple-sector real business cycle model introduces not unmanageable complexity, but qualitatively new behavior that in some cases can be quite manageable. Our findings also show that calibrations aimed at matching only a few first and second moments can similarly hide major differences between models and the data, missing long-run dependence properties. While widening the theoretical horizons of the paradigm, fractional techniques also widen the potential testing of such theories.

3.2. Fiscal Policy and Welfare Implications

Taking a policy perspective raises two natural questions about the fractional properties of national income. First, will fiscal or monetary policy change the degree of long-term dependence? Friedman and Schwartz (1982), for example, point out that long-run income cycles correlate with long-run monetary cycles. Second, does long-term dependence have welfare implications? Do agents care that they live in such a world?

In the basic Ramsey-Solow growth model, as in its stochastic extensions, taxes affect output and capital levels but not growth rates; thus, tax policy does not affect fractional properties.[14] However, two alternative approaches suggest richer possibilities. First, recall that fractional noise arises through the aggregation of many autoregressive processes. Fiscal policy may not change the coefficients of each process, but a tax policy can alter the distribution of total output across individuals, effectively changing the fractional properties of the aggregate. Second, endogenous growth models often allow tax policy to affect growth rates by reducing investment in research, thus depressing future growth.[15] Hence, the autoregressive parameters of an individual firm's output could change with policy, in turn affecting aggregate income.

Unfortunately, implementing either approach with even a modicum of realism would be quite complicated. In the dynamic stochastic growth model, taxation drives a wedge between private and social returns, resulting in a suboptimal equilibrium. This eliminates methods that exploit the pareto-optimality of competitive equilibrium, such as dynamic programming. Characterizing solutions requires simulation methods, because no closed forms have been found.[16] Thus, it seems clear that fiscal policy can affect fractional properties. Explicitly calculating the impact would take this paper too far afield and is best left for future research.

Those who forecast output or sales will care about the fractional nature of output, but fractional processes can have normative implications as well. Following Lucas (1987), this section estimates the welfare costs of economic instability under different regimes. We can decide if people care whether their world is fractional. For concreteness, let the typical household consume C_t, evaluating this via a utility function:

$$U = E\left[\sum_{t=0}^{\infty} \beta^t \frac{C^{1}_{t-\sigma}}{1-\sigma}\right] = \sum_{t=0}^{\infty} \frac{\beta^t}{1-\sigma} E[C_t^{1-\sigma}]. \tag{3.21}$$

Also assume that

$$\ln C_t = (1+\lambda) \sum_{k=0}^{\infty} \phi_k L^k \eta_t, \tag{3.22}$$

where $\eta_t \equiv \ln \epsilon_t$. The λ term measures compensation for variations in the process $\phi(L)$. With η_t normally distributed with mean zero and variance one, the compensating fraction λ between two processes ϕ and ψ is

$$1 + \lambda = \exp \left[\frac{1}{2} (1 - \sigma) \sum_{k=0}^{\infty} (\psi_k^2 - \phi_k^2) \right]. \tag{3.23}$$

Evaluating (3.23) using a realistic $\sigma = 5$, again comparing an AR(1) with $\rho =$ 0.9 against a fractional process of order one-fourth, we find that $\lambda =$ -0.99996.(This number looks larger than those in Lucas [1987] because the process is in logs rather than in levels.[17]) For comparison, this is the difference between an AR(1) with ρ of 0.90 and one with ρ of 0.95. This calculation provides only a rough comparison. When feasible, welfare calculations should use the model generating the processes, as only it will correctly account for important specifics such as labor supply or distortionary taxation.

4. Rescaled Range Analysis of Real Output

The results in section 3 show that simple aggregation may be one source of long-term dependence in the business cycle. In this section, we employ a method for detecting long memory and apply it to real GNP. The technique is based on a simple generalization of a statistic first proposed by the English hydrologist Harold Edwin Hurst (1951) and subsequently refined by Mandelbrot (1972, 1975) and others.[18] Our generalization of Mandelbrot's statistic, called the rescaled range, the range over standard deviation, or the R/S statistic, enables us to distinguish between short- and long-run dependence, in a sense that will be made precise below. We define our notions of short and long memory and present the test statistic in section 4.1. Section 4.2 gives the empirical results for real GNP. We find long-term dependence in log-linearly detrended output, but considerably less dependence in the growth rates. To interpret these findings, we perform several Monte Carlo experiments under two null and two alternative hypotheses. Results are reported in section 4.3.

4.1. The R/S Statistic

To develop a method of detecting long memory, we must be precise about the distinction between long-term and short-term statistical dependence. One of the most widely used concepts of short-term dependence is the notion of "strong-mixing" (based on Rosenblatt [1956]), a measure of the decline in statistical dependence of two events separated by successively longer time spans. Heuristically, a time series is strong-mixing if the maximal dependence between any two events becomes trivial as more time elapses between them. By controlling the rate at which the dependence between future events and those of the distant past declines, it is possible to extend the usual laws of large numbers and central-limit theorems to dependent sequences of random variables. Such mixing conditions have been used extensively by White (1980), White and Domowitz (1984), and Phillips (1987), for example, to relax the assumptions that ensure consistency and asymptotic normality of various econometric estimators. We adopt this notion of short-term dependence as part of our null hypothesis. As Phillips (1987) observes, these conditions are satisfied by a great many stochastic processes, including all Gaussian finite-order stationary ARMA models. Moreover, the inclusion of a moment condition allows for heterogeneously distributed sequences (such as those exhibiting heteroscedasticity), an especially important extension in view of the nonstationarities of real GNP.

In contrast to the "short memory" of weakly dependent (i.e., strong-mixing) processes, natural phenomena often display long-term memory in the form of nonperiodic cycles. This has led several authors, most notably Mandelbrot, to develop stochastic models that exhibit dependence even over very long time spans. The fractionally integrated time-series models of Mandelbrot and Van Ness (1968), Granger and Joyeux (1980), and Hosking (1981) are examples of these. Operationally, such models possess autocorrelation functions that decay at much slower rates than those of weakly dependent processes, violating the conditions of strong-mixing. To detect long-term dependence (also called strong dependence), Mandelbrot suggests using the R/S statistic, which is the range of partial sums of deviations of a time series from its mean, rescaled by its standard deviation. In several seminal papers, Mandelbrot demonstrates the superiority of the R/S statistic over more conventional methods of determining long-run dependence, such as autocorrelation analysis and spectral analysis.[19]

In testing for long memory in output, we employ a modification of the R/S statistic that is robust to weak dependence. In Lo (1991), a formal sampling theory for the statistic is obtained by deriving its limiting distribution analytically using a functional central-limit theorem.[20] We use this statistic and its asymptotic distribution for inference below. Let X_t denote the first difference of log-GNP; we assume that

$$X_t = \mu + \epsilon_t, \tag{4.1}$$

where μ is an arbitrary but fixed parameter. Whether or not X_t exhibits long-term memory depends on the properties of ϵ_t. For the null hypothesis H, the sequence of disturbances ϵ_t satisfies the following conditions:

(A1) $E[\epsilon_t] = 0$ for all t.

(A2) $\sup_t E\left[\,|\epsilon_t|^{\beta}\right] < \infty$ for some $\beta > 2$.

(A3) $\sigma^2 = \lim_{n\to\infty} E\left[\frac{1}{n}\left(\sum_j^n\right)^2\right]$ exists, and $\sigma^2 > 0$.

(A4) $\{\epsilon_t\}$ is strong-mixing, with mixing coefficients α_k that satisfy[21]

$$\sum_{k=1}^{\infty} \alpha_k^{1-\frac{2}{\beta}} < \infty.$$

Condition (A1) is standard. Conditions (A2) through (A4) are restrictions on the maximal degree of dependence and heterogeneity allowable while still permitting some form of the law of large numbers and the (functional) central-limit theorem to obtain. Note that we have not assumed stationarity. Although condition (A2) rules out infinite-variance marginal distributions of ϵ_t, such as those in the stable family with characteristic exponent less than two, the disturbances may still exhibit leptokurtosis via time-varying conditional moments (e.g., conditional heteroscedasticity). Moreover, since there is a trade-off between conditions (A2) and (A4), the uniform bound on the moments may be relaxed if the mixing coefficients decline faster than (A4) requires.[22] For example, if we require ϵ_t to have finite absolute moments of all orders (corresponding to $\beta \to \infty$), then α_k must decline faster than 1/k. However, if we restrict ϵ_t to have finite moments only up to order four, then α_k must decline faster than $1/k^2$. These conditions are discussed at greater length in Phillips (1987), to which we refer interested readers.

Conditions (A1) through (A4) are satisfied by many of the recently proposed stochastic models of persistence, such as the stationary AR(1) with a near-unit root. Although the distinction between dependence in the short versus the long run may appear to be a matter of degree, strongly dependent processes behave so differently from weakly dependent ones that our dichotomy seems quite natural. For example, the spectral densities of strongly dependent processes are either unbounded or zero at frequency zero. Their partial sums do not converge in distribution at the same rate as weakly dependent series, and graphically, their behavior is marked by cyclic patterns of all kinds, some that are virtually indistinguishable from trends.[23]

To construct the modified R/S statistic, consider a sample X_1, X_2, ..., X_n and let $\bar{X}_n$ denote the sample mean $\frac{1}{n} \sum_j X_j$. Then, the modified R/S statistic, which we shall call Q_n, is given by

$$Q_n \equiv \frac{1}{\hat{\sigma}_n(q)} \left[\underset{1\leq k\leq n}{\operatorname{Max}} \sum_{j=1}^{k} \left(X_j - \bar{X}_n \right) - \underset{1\leq k\leq n}{\operatorname{Min}} \sum_{j=1}^{k} \left(X_j - \bar{X}_n \right) \right], \tag{4.2}$$

where

$$\hat{\sigma}_n^2(q) \equiv \frac{1}{n} \sum_{j=1}^{n} (X_j - \bar{X}_n) \Big]^2 + \frac{2}{n} \omega_j(q) \left\{ \sum_{i=j+1}^{n} (X_i - \bar{X}_n)(X_{i-j} - \bar{X}_n) \right\} \tag{4.3}$$
$$= \hat{\sigma}_x^2 + 2 \sum_{j=1}^{q} \omega_j(q)\, \hat{\gamma}_j \qquad \omega_j(1) \equiv 1 - \frac{j}{q+1}\; q < n,$$

and $\hat{\sigma}_x^2$ and $\hat{\gamma}_j$ are the usual sample variance and autocovariance estimators of X. Q_n is the range of partial sums of deviations of X_j from its mean, $\bar{X}_n$, normalized by an estimator of the partial sum's standard deviation divided by n. The estimator $\hat{\sigma}_n(q)$ involves not only sums of squared deviations of X_j, but also its weighted autocovariances up to lag q; the weights $\omega_j(q)$ are those suggested by Newey and West (1987), and they always yield a positive estimator $\hat{\sigma}_n^2(q)$.[24] Theorem 4.2 in Phillips (1987) demonstrates the consistency of $\hat{\sigma}_n(q)$ under the following conditions:

(A2′) $\sup_t E[|\epsilon_t|^{2\beta}] < \infty$ for some $\beta > 2$.

(A5) As n increases without bound, q also rises without bound, such that $q \sim o(n^{1/4})$.

The choice of the truncation lag q is a delicate matter. Although q must increase with the sample size (although at a slower rate), Monte Carlo evidence suggests that when q becomes large relative to the number of observations, asymptotic approximations may fail dramatically.[25] If the chosen q is too small, however, the effects of higher-order autocorrelations may not be captured. Clearly, the choice of q is an empirical issue that muust take into account the data at hand.

Under conditions (A1), (A2′), (A3) ... A(5), Lo (1991) shows that the statistic $V_n \equiv Q_n/\sqrt{n}$ has a well-defined asymptotic distribution given by the random variable V, whose distribution function $F_V(\nu)$ is[26]

$$F_V(\nu) = \sqrt{2\pi} \sum_{k=-\infty}^{\infty} \left\{ k[(1-\alpha_k^2)\phi(\alpha_k) - k\phi(\beta_k) \right\} \tag{4.4}$$

$$\alpha_k \equiv 2k\nu \quad \beta_k \equiv 2(k+1)\nu \quad \phi(x) \equiv \frac{1}{\sqrt{2\pi}} e^{-\frac{1}{2}x^2} .$$

Using F_v, critical values may be readily calculated for tests of any significance level. The most commonly used values are reported in tables la and lb. Table la reports the fractiles of the distribution, while table lb reports the symmetric confidence intervals about the mean. The moments of V are also easily computed using the density function f_V; it is straightforward to show that $E[V] = \sqrt{\frac{\pi}{2}}$ and $E[V^2] = \frac{\pi^2}{6}$. Thus, the mean and standard deviation of V are approximately 1.25 and 0.27, respectively. The distribution and density functions are plotted in figure 4. Note that the distribution is positively skewed and that most of its mass falls between three-fourths and two.

If the observations are independently and identically distributed with variance σ_ϵ^2, our normalization by $\hat{\sigma}_n(q)$ is asymptotically equivalent to normalizing by the usual standard deviation estimator $s_n = [\frac{1}{n} \sum_j (X_j - \bar{X}_n)^2]^{1/2}$. The resulting statistic, which we call $\tilde{Q}_n$, is precisely the one proposed by Hurst (1951) and Mandelbrot (1972):

$$\tilde{Q}_n \equiv \frac{1}{s_n} \left[\underset{1\leq k\leq n}{\mathrm{Max}} \sum_{j=1}^{k} (X_j - \bar{X}_n) - \underset{1\leq k\leq n}{\mathrm{Min}} \sum_{j=1}^{k} (X_j - \bar{X}_n) \right] . \tag{4.5}$$

Under the more restrictive null hypothesis of i.i.d. observations, the statistic $\tilde{V}_n \equiv \tilde{Q}_n/\sqrt{n}$ can be shown to converge to V as well. However, in the presence of short-range dependence, $\tilde{V}_n$ does not converge to V, whereas V_n still does. Of course, if the particular form of short-range dependence is known, it can be accounted for in deriving the limiting distribution of $\tilde{V}_n$. For example, if X_t is a stationary AR(1) with autoregressive parameter ρ, Lo (1991) shows that $\tilde{V}_n$ converges to ξV, where $\xi \equiv \sqrt{(1+\rho)/(1-\rho)}$. But since we would like our limiting distribution to be robust to *general* forms of short-range dependence, we use the modified R/S statistic V_n below.

4.2. Empirical Results for Real Output

We apply our test to two time series of real output: quarterly postwar real GNP from 1947:IQ to 1987:IVQ, and the annual Friedman and Schwartz (1982) series from 1869 to 1972. These results are reported in table 2. Entries in the first numerical row are estimates of the classical R/S statistic $\tilde{V}_n$, which is not robust to short-term dependence. The next eight rows are estimates of the modified R/S statistic $V_n(q)$ for values of q from one to eight. Recall that q is the truncation lag of the spectral density estimator at frequency zero. Reported in parentheses below the entries for $V_n(q)$ are estimates of the percentage bias of the statistic $\tilde{V}_n$, computed as $100 \cdot [\tilde{V}_n/V_n(q) - 1]$.

The first column of numerical entries in table 2 indicates that the null hypothesis of short-term dependence for the first difference of log-GNP cannot be rejected for any value of q. The classical R/S statistic also supports the null hypothesis, as do the results for the Friedman and Schwartz series. On the other hand, when we log-linearly detrend real GNP, the results are considerably different. The third column of numerical entries in table 2 shows that short-term dependence may be rejected for log-linearly detrended quarterly output with values of q from one to four. That the rejections are weaker for larger q is not surprising, since additional noise arises from estimating higher-order autocorrelations. When values of q beyond four are used, we no longer reject the null hypothesis at the 5 percent level of significance. Finally, using the Friedman and Schwartz time series, we only reject with the classical R/S statistic and with $V_n(1)$.

The values reported in table 2 are qualitatively consistent with the results of other empirical studies of fractional processes in GNP, such as Diebold and Rudebusch (1989) and Sowell (1989). For first differences, the R/S statistic falls below the mean, suggesting a negative fractional exponent, or in level terms, an exponent between zero and one. Furthermore, though the earlier papers produce point estimates, the imprecision of these estimates means that they do not reject the hypothesis of short-term dependence. For example, the standard-deviation error bounds for Diebold and Rudebusch's two point estimates, $d = 0.9$ and $d = 0.52$, are (0.42, 1.38) and (0.06, 1.10), respectively.

Taken together, our results confirm the unit-root findings of Campbell and Mankiw (1987), Nelson and Plosser (1982), Perron and Phillips (1987), and Stock and Watson (1986). That there are more significant autocorrelations in log-linearly detrended GNP is precisely the spurious periodicity suggested by Nelson and Kang (1981). Moreover, the trend plus stationary noise model of GNP is not contained in our null hypothesis; hence, our failure to reject the null hypothesis is also consistent with the unit-root model.[27] To see this, observe that if log-GNP y_t were trend stationary, i.e., if

$$y_t = \alpha + \beta t + \eta_t, \tag{4.6}$$

where η_t is stationary white noise, then its first difference X_t would simply be $X_t = \beta + \epsilon_t$, where $\epsilon_t \equiv \eta_t - \eta_{t-1}$. But this innovations process violates our assumption (A3) and is therefore not contained in our null hypothesis.

Sowell (1989) has used estimates of d to argue that the trend-stationary model is correct. Following the lead of Nelson and Plosser (1982), he investigates whether the d parameter for the first-differenced series is close to zero, as the unit-root specification suggests, or close to minus one, as the trend-stationary specification suggests. His estimate of d is in the general range of -0.9 to -0.5, providing some evidence that the trend-stationary interpretation is correct. Even in this case, however, the standard errors tend to be large, on the order of 0.36. Although our procedure yields no point estimate of d, it does seem to rule out the trend-stationary case.

To conclude that the data support the null hypothesis because our statistic fails to reject it is premature, of course, since the size and power of our test in finite samples is yet to be determined.

4.3. Size and Power of the Test

To evaluate the size and power of our test in finite samples, we perform several illustrative Monte Carlo experiments for a sample of 163 observations, which corresponds to the number of quarterly observations of real GNP growth from 1947:IQ to 1987:IVQ.[28] We simulate two null hypotheses:

independently and identically distributed increments, and increments that follow an ARMA(2,2) process. Under the i.d.d. null hypothesis, we fix the mean and standard deviation of our random deviates to match the sample mean and standard deviation of our quarterly data set: 7.9775×10^{-3} and 1.0937×10^{-3}, respectively. To choose parameter values for the ARMA(2,2) simulation, we estimate the model

$$(1- \phi_1 L - \phi_2 L^2) y_t = \mu + (1 + \theta_1 L + \theta_2 L^2)\epsilon_t \qquad \epsilon_t \sim WN(0,\sigma_\epsilon^2) \tag{4.7}$$

using nonlinear least squares. The parameter estimates are as follows (standard errors are in parentheses):

$$\phi_1 = 0.5837 \qquad \theta_1 = -0.2825$$
$$(0.1949) \qquad (0.1736)$$
$$\phi_2 = -0.4844 \qquad \theta_2 = 0.6518$$
$$(0.1623) \qquad (0.1162)$$
$$\mu = 0.0072$$
$$(0.0016)$$
$$\hat{\sigma}_\epsilon^2 = 0.0102$$

Table 3 reports the results of both null simulations.

It is apparent from the i.i.d. null panel of table 3 that the 5 percent test based on the classical R/S statistic rejects too frequently. The 5 percent test using the modified R/S statistic with q = 3 rejects 4.6 percent of the time, closer to the nominal size. As the number of lags increases to eight, the test becomes more conservative. Under the ARMA(2,2) null hypothesis, it is apparent that modifying the R/S statistic by the spectral density estimator $\hat{\sigma}_n^2(q)$ is critical. The size of a 5 percent test based on the classical R/S statistic is 34 percent, whereas the corresponding size using the modified R/S statistic with q = 5 is 4.8 percent. As before, the test becomes more conservative when q is increased.

Table 3 also reports the size of tests using the modified R/S statistic when the lag length q is optimally chosen using Andrews' (1987) procedure. This data-dependent procedure entails computing the first-order autocorrelation coefficient $\hat{\rho}(1)$ and then setting the lag length as the integer value of $\bar{M}_n$, where[29]

$$\bar{M}_n \equiv \left(\frac{3\hat{\alpha}n}{2}\right)^{1/3} \quad \hat{\alpha} \equiv \frac{4\hat{\rho}^2}{(1-\hat{\rho}^2)^2}. \tag{4.8}$$

Under the i.i.d. null hypothesis, Andrews' formula yields a 5 percent test with empirical size 6.9 percent; under the ARMA(2,2) alternative, the corresponding figure is 4.1 percent. Although significantly different from the nominal value, the empirical size of tests based on Andrews' formula may not be economically important. In addition to its optimality properties, the procedure has the advantage of eliminating a dimension of arbitrariness from the test. Table 4 reports power simulations under two fractionally differenced alternatives: $(1-L)^d\epsilon_t = \eta_t$, where d = (1/3, -1/3). Hosking (1981) has shown that the autocovariance function $\gamma_\epsilon(k)$ equals

$$\frac{\Gamma(1 - 2d)\Gamma(d + k)}{\Gamma(d)\Gamma(1 - d)\Gamma(1 - d + k)}\sigma_\eta^2 \quad d \in (-\tfrac{1}{2}, \tfrac{1}{2}). \tag{4.9}$$

Realizations of fractionally differenced time series of length 163 are simulated by pre-multiplying vectors of independent standard normal random variates by the Cholesky factorization of the 163 x 163 covariance matrix, whose entries are given by (3.11). To calibrate the simulations, σ_η^2 is chosen to yield unit variance ϵ_t. We then multiply the ϵ_t series by the sample standard deviation of real GNP growth from 1947:IQ to 1987:IVQ and add the sample mean of real GNP growth over the same period. The resulting time series is used to compute the power of the R/S statistic (see table 4).

For small values of q, tests based on the modified R/S statistic have reasonable power against both of the fractionally differenced alternatives. For example, using one lag, the 5 percent test has 58.7 percent power against the d = 1/3 alternative and 81.1 percent power against the d = -1/3 alternative. As the lag is increased, the test's power declines.

Note that tests based on the classical R/S statistic are significantly more powerful than those using the modified R/S statistic. This, however, is of little value when distinguishing between long-term and short-term dependence, since the test using the classical statistic also has power against some stationary finite-order ARMA processes. Finally, note that tests using Andrews' truncation lag formula have reasonable power against the d = -1/3 alternative, but are considerably weaker against the more relevant d = 1/3 alternative.

The simulation evidence in tables 3 and 4 suggests that our empirical results do indeed support the short-term dependence of GNP with a unit root. Our failure to reject the null hypothesis does not seem to be explicable by a lack of power against long-memory alternatives. Of course, our simulations are illustrative and by no means exhaustive; additional Monte Carlo experiments will be required before a full assessment of the test's size and power is complete. Nevertheless, our modest simulations indicate that there is little empirical evidence of long-term memory in GNP growth rates. Perhaps a direct estimation of long-memory models would yield stronger results, an issue that has recently been investigated by several authors.[30]

5. Conclusion

This paper has suggested a new approach for investigating the stochastic structure of aggregate output. Traditional dissatisfaction with conventional methods -- from observations about the typical spectral shape of economic time series to the discovery of cycles at all periods -- calls for such a reformation. Indeed, recent controversy about deterministic versus stochastic trends and the persistence of shocks underscores the difficulties even modern methods have in identifying the long-run properties of the data.

Fractionally integrated random processes provide one explicit approach to the problem of long-term dependence; naming and characterizing this aspect is the first step in studying the problem scientifically. Controlling for long-term dependence improves our ability to isolate business cycles from trends and to assess the propriety of that decomposition. To the extent that long-term dependence explains output, it deserves study in its own right. Furthermore, Singleton (1988) has pointed out that dynamic macroeconomic models often inextricably link predictions about business cycles, trends, and seasonal effects. So, too, is long-term dependence linked: A fractionally integrated process arises quite naturally in a dynamic linear model via aggregation. Our model not only predicts the existence of fractional noise, but suggests the character of its parameters. This class of models leads to testable restrictions on the nature of long-term dependence in aggregate data, and also holds the promise of enhancing policy evaluation.

Advocating a new class of stochastic processes would be a fruitless task if its members were intractable. But in fact, manipulating such processes causes few problems. We construct an optimizing linear dynamic model that exhibits fractionally integrated noise, and provide an explicit test for such long-term dependence. Modifying a statistic developed by Hurst and Mandelbrot gives us a statistic robust to short-term dependence. This modified R/S statistic possesses a well-defined limiting distribution, which we have tabulated. Illustrative computer simulations indicate that this test has power against at least two specific alternative hypotheses of long-term memory.

Two main conclusions arise from our empirical work and from Monte Carlo experiments. First, the evidence does not support long-term dependence in GNP. Rejections of the short-term-dependence null hypothesis occur only with detrended data and are consistent with the well-known problem of spurious periodicities induced by log-linear detrending. Second, since a trend-stationary model is not contained in our null hypothesis, our failure to reject may also be viewed as supporting the first-difference stationary model of GNP, with the additional result that the stationary process is at best weakly dependent. This supports and extends Adelman's conclusion that, at least within the confines of the available data, there is little evidence of long-term dependence in the business cycle.

Footnotes

1. The idea of fractional differentiation is an old one (dating back to an oblique reference by Leibniz in 1695), but the subject lay dormant until the nineteenth century, when Abel, Liouville, and Riemann developed it more fully. Extensive applications have only arisen in this century; see, for example, Oldham and Spanier (1974). Kolmogorov (1940) was apparently the first to notice its applicability in probability and statistics.

2. When d is an integer, (2.3) reduces to the better-known formula for the binomial coefficient, $\frac{d!}{k!(d-k)!}$. We follow the convention that $\binom{d}{0} = 1$ and $\binom{0}{0} = 0$.

3. See Hosking (1981) for further details.

4. See Cochrane (1988) and Quah (1987) for opposing views.

5. There has been some confusion about this point in the literature. Geweke and Porter-Hudak (1983) argue that $C(1) > 0$. They correctly point out that Granger and Joyeux (1980) erred, but then incorrectly claim that $C(1) = 1/\Gamma(d)$. If our equation (2.7) is correct, then it is apparent that $C(1) = 0$ (which agrees with Granger [1980] and Hosking [1981]). Therefore, the focus of the conflict lies in the approximation of the ratio $\Gamma(k+d)/\Gamma(k+1)$ for large k. We have used Stirling's approximation. However, a more elegant derivation follows from the functional analytic definition of the gamma function as the solution to the following recursive relation (see, for example, Iyanaga and Kawada [1980, section 179.A]):
$$\Gamma(x+1) = x\Gamma(x)$$
and the conditions
$$\Gamma(1) = 1 \quad \lim_{n\to\infty} \frac{\Gamma(x+n)}{n^x\Gamma(n)} = 1.$$

6. See Chatfield (1984, chapters 6 and 9).

7. See Sargent (1987, chapter 1) for an excellent exposition.

8. See Theil (1954).

9. Granger (1980) conjectures that this particular distribution is not essential.

10. For a discussion of the variety of shapes the beta distribution can take as p and q vary, see Johnson and Kotz (1970).

11. Two additional points are worth emphasizing. First, the beta distribution need not be over (0,1) to obtain these results, only over $(\alpha,1)$. Second, it is indeed possible to vary the a_i's so that α_i has a beta distribution.

12. Leontief, in his classic (1976) study, reports own-industry output coefficients for 10 sectors, investigating how much an extra unit of food will increase food production. Results vary from 0.06 (fuel) to 1.24 (other industries).

13. See Jorgenson, Gollop, and Fraumeni (1987).

14. See Atkinson and Stiglitz (1980).

15. For example, see Romer (1986) and King, Plosser, and Rebelo (1987).

16. See King, Plosser, and Rebelo (1987), Baxter (1988), and Greenwood and Huffman (1991).

17. We calculate this using (2.7) and the Hardy-Littlewood approximation for the resulting Rieman Zeta Function, following Titchmarsh (1951, section 4.11).

18. See Mandelbrot and Taqqu (1979) and Mandelbrot and Wallis (1968, 1969a-c).

19. See Mandelbrot (1972, 1975), Mandelbrot and Taqqu (1979), and Mandelbrot and Wallis (1968, 1969a-c).

20. This statistic is asymptotically equivalent to Mandelbrot's under independently and identically distributed observations. However, Lo (1991) shows that the original R/S statistic may be significantly biased toward rejection when the time series is short-term dependent. Although aware of this bias, Mandelbrot (1972, 1975) did not correct for it, since his focus was on the relation of the R/S statistic's logarithm to the logarithm of the sample size, which involves no statistical inference; such a relation clearly is unaffected by short-term dependence.

21. Let $\{\epsilon_t(\omega)\}$ be a stochastic process on the probability space (Ω, F, P) and define

$$\alpha(A,B) \equiv \sup_{\{A\in A,\, B\in B\}} |P(A\cap B) - P(A)P(B)| \quad A\subset F, B\subset F$$

The quantity $\alpha(A,B)$ is a measure of the dependence between the two σ fields A and B in F. Denote by B_s^t the Borel σ field generated by $[\epsilon_t(\omega), \ldots, \epsilon_t(\omega)]$, i.e., $B_s^t \equiv \sigma[\epsilon_t(\omega), \ldots, \epsilon_t(\omega)] \subset F$. Define the coefficients α_k as

$$\alpha_k \equiv \sup_j \alpha\,(B^j{}_{-\infty},\ B_{j+k}^{\infty}).$$

Then, $\{\epsilon_t(\omega)\}$ is said to be strong-mixing if $\lim_{G\to\infty} \alpha_k = 0$. For further details, see Rosenblatt (1956), White (1984), and the papers in Eberlein and Taqqu (1986).

22. See Herndorf (1985). Note that one of Mandelbrot's (1972) arguments in favor of R/S analysis is that finite second moments are not required. This is indeed the case if we are interested only in the almost sure convergence of the statistic. However, since we wish to derive its *limiting distribution* for purposes of inference, a stronger moment condition is needed.

23. See Mandelbrot (1972) for further details.

24. $\sigma_n^2(q)$ is also an estimator of the spectral density function of X_t at frequency zero, using a Bartlett window.

25. See, for example, Lo and MacKinlay (1988).

26. V may be shown to be the range of a Brownian bridge on the unit interval. See Lo (1991) for further details.

27. Of course, this may be the result of low power against stationary but near-integrated processes, an issue that must be addressed by Monte Carlo experiments.

28. All simulations were performed in double precision on a VAX 8700 using the IMSL 10.0 random number generator DRNNOA. Each experiment consisted of 10,000 replications.

29. In addition, Andrews' procedure requires weighting the autocovariances by $1 - \frac{j}{M_n}$ $(j = 1, \ldots, [\bar{M}_n])$, in contrast to Newey and West's (1987) $1 - \frac{j}{q+1}$ $(j = 1, \ldots, q)$, where q is an integer but $(\bar{M}_n)$ need not be.

30. See, for example, Diebold and Rudebusch (1989), Sowell (1987), and Yajima (1985, 1988).

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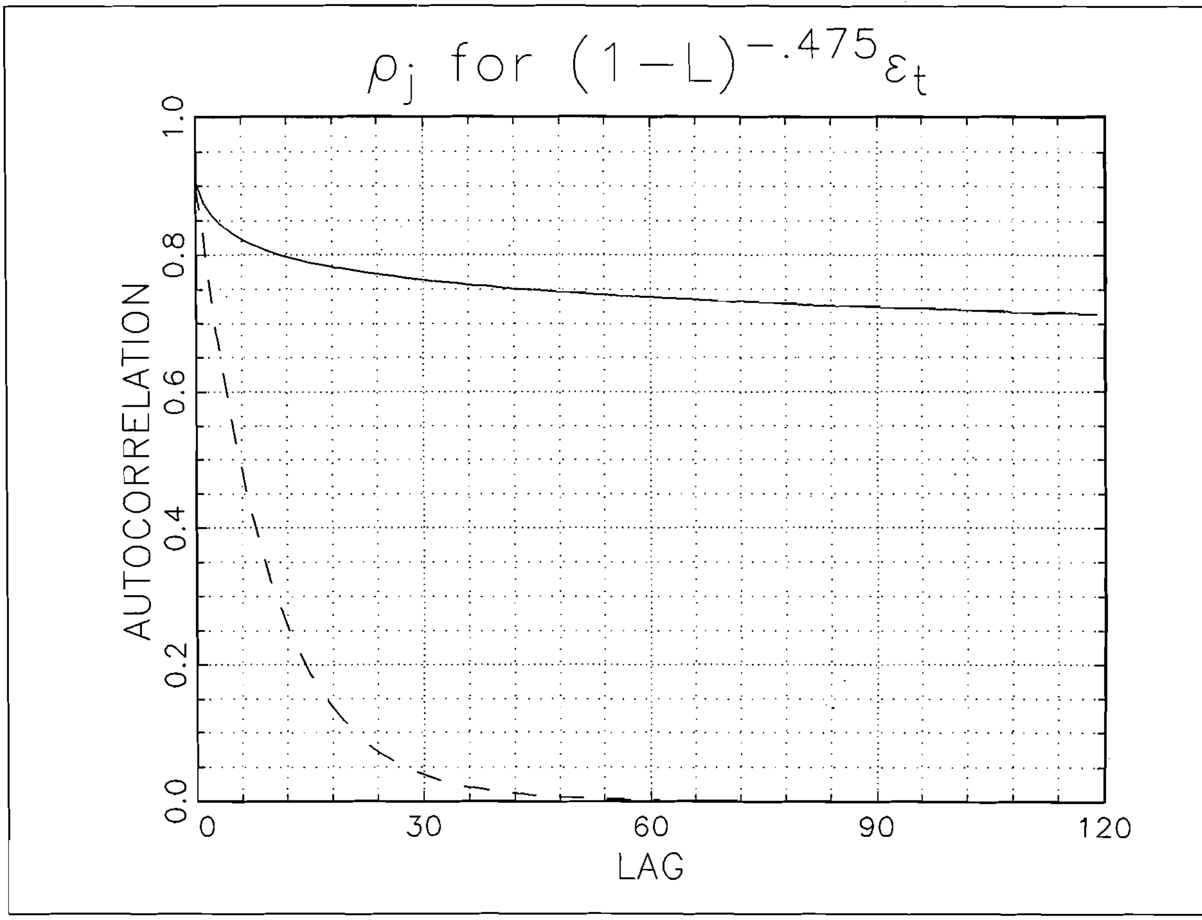

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Figure 1

Autocorrelation functions of an AR(1) with coefficient 0.90 (dashed line) and a fractionally differenced series $X_t = (1 - L)^{-d}\epsilon_t$ with differencing parameter $d = 0.475$ (solid line). Although both processes have a first-order autocorrelation of 0.90, the fractionally differenced process decays much more slowly.

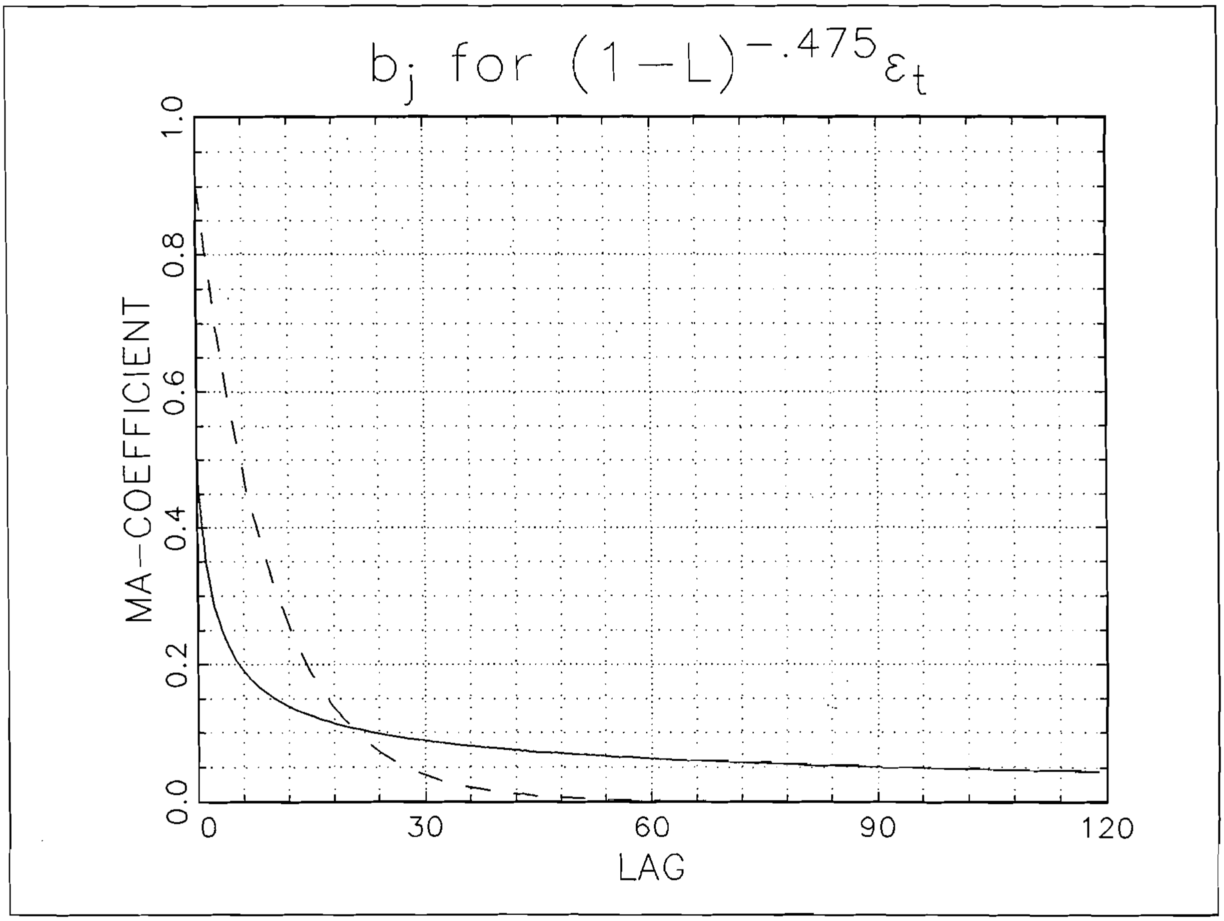

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Figure 2

Impulse-response function (solid line) of the fractionally differenced time series $X_t = (1 - L)^{-d}\epsilon_t$ for differencing parameter d = 0.475. For comparison, the impulse-response function of an AR(1) with autoregressive parameter 0.90 is also plotted (dashed line).

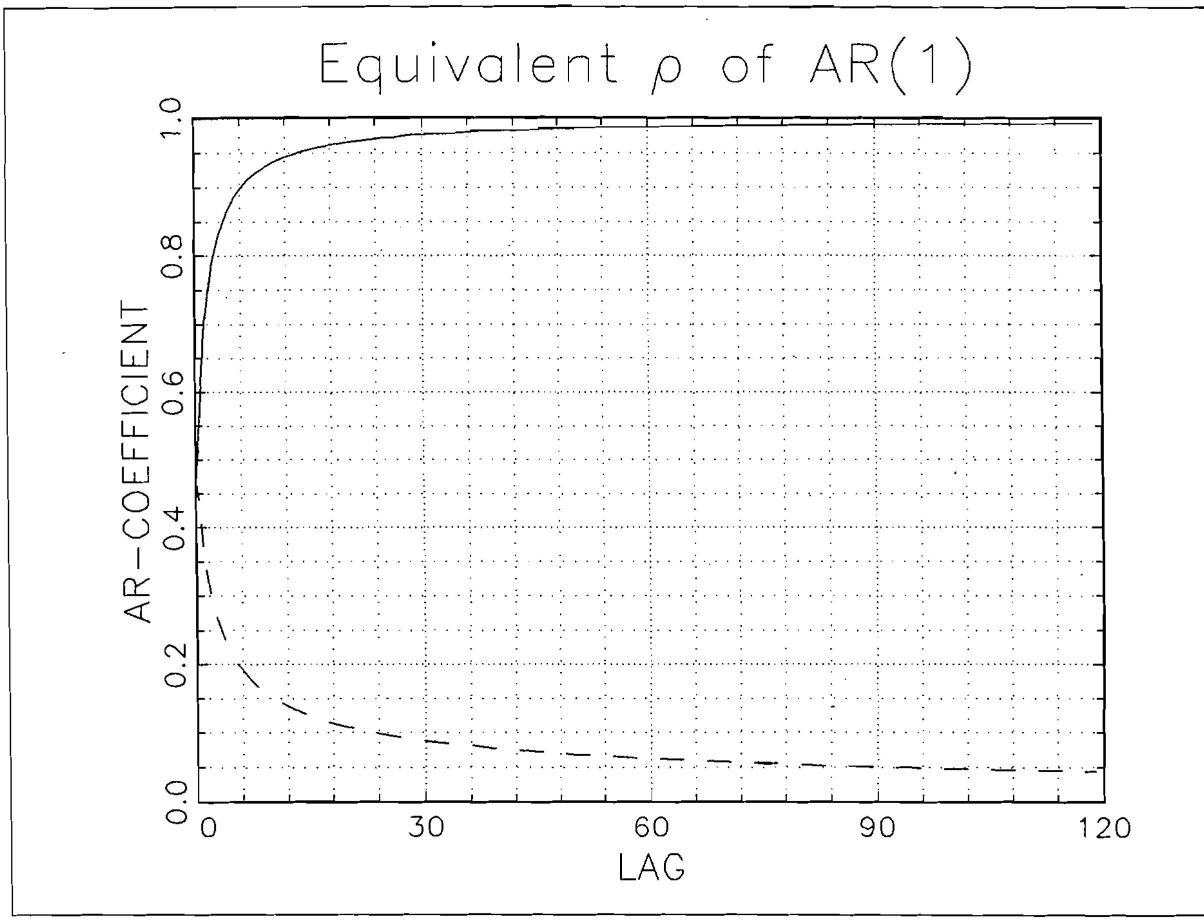

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Figure 3

Values of an AR(1)'s autoregressive parameter required to generate the same k-th order autocorrelation as the fractionally differenced series $X_t = (1 - L)^{-d}\epsilon_t$ for differencing parameter d = 0.475 (solid line). Formally, this is simply the k-th root of the fractionally differenced series' impulse-response function (dashed line). For large k, the autoregressive parameter must be very close to unity.

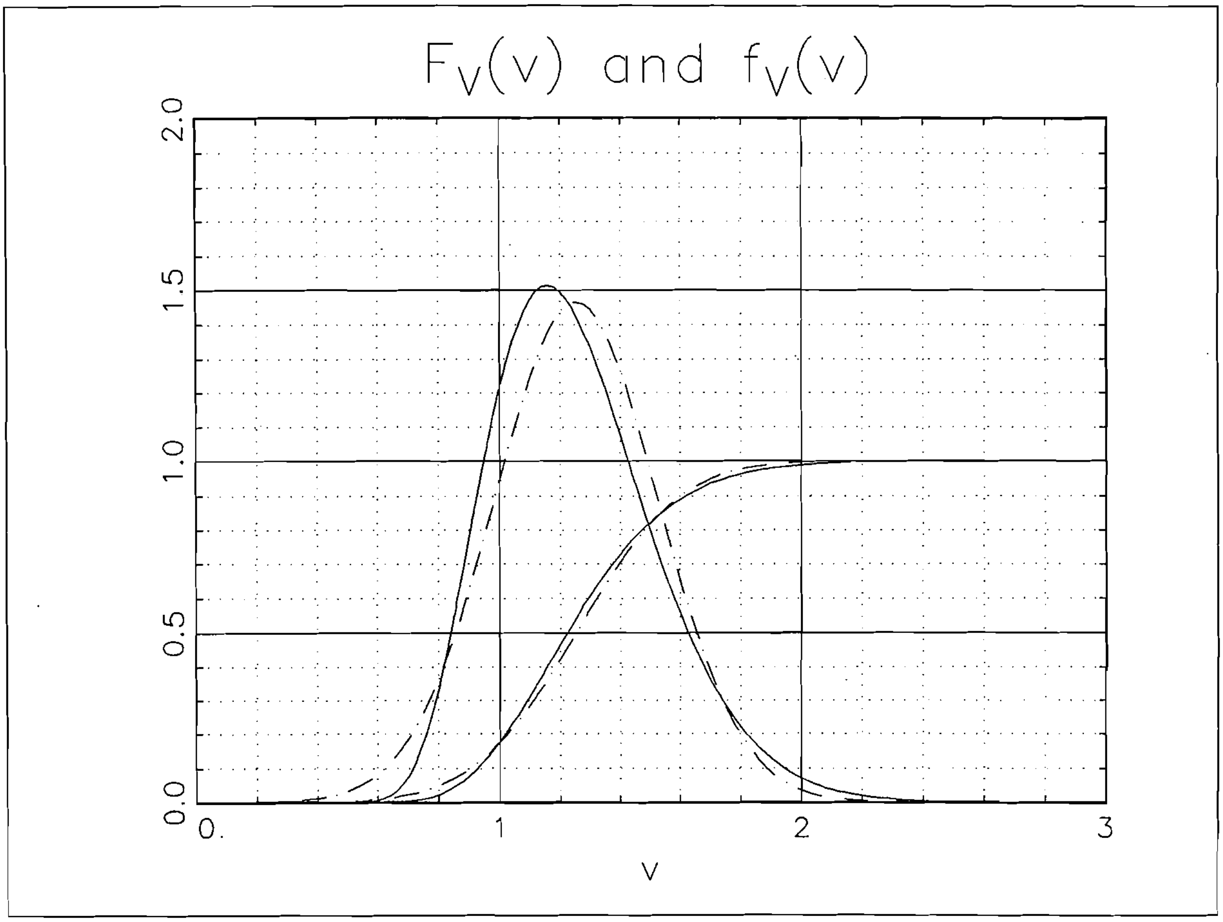

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Figure 4

Distribution and density function of the range V of a Brownian bridge. Dashed curves are the normal distribution and density functions with mean and variance equal to those of V.

Table 1a. Fractiles of the Distribution $F_V(v)$

$P(V < v)$	.005	.025	.050	.100	.200	.300	.400	.500
v	0.721	0.809	0.861	0.927	1.018	1.090	1.157	1.223

$P(V < v)$	.543	.600	.700	.800	.900	.950	.975	.995
v	$\sqrt{\frac{\pi}{2}}$	1.294	1.374	1.473	1.620	1.747	1.862	2.098

Table 1b. Symmetric Confidence Intervals about the Mean

$P(\sqrt{\frac{\pi}{2}} - \gamma < V < \sqrt{\frac{\pi}{2}} + \gamma)$	γ
.001	0.748
.050	0.519
.100	0.432
.500	0.185

Source: Authors

Table 2

R/S analysis of real GNP; $y^{Q}_{Det.}$ indicates log-linearly detrended quarterly real GNP from 1947:IQ to 1987:IVQ, and Δy^{Q} indicates the first differences of the logarithm of real GNP. $y^{FS}_{Det.}$ and Δy^{FS} are defined similarly for the Friedman and Schwartz series. The classical R/S statistic $\tilde{V}_n$ and the modified R/S statistic $V_n(q)$ are reported.[1]

	Δy^{Q}	Δy^{FS}	$y^{Q}_{Det.}$	$y^{FS}_{Det.}$
$\tilde{V}_n$	1.25	1.00	4.23*	2.83*
$V_n(1)$	1.07	0.94	3.02*	2.10*
(%–Bias)	(17.2)	(6.6)	(40.0)	(35.2)
$V_n(2)$	0.97	0.93	2.49*	1.79
(%–Bias)	(29.0)	(7.3)	(69.6)	(58.5)
$V_n(3)$	0.93	0.95	2.19*	1.62
(%–Bias)	(34.6)	(4.7)	(93.5)	(74.9)
$V_n(4)$	0.92	1.00	1.98*	1.52
(%–Bias)	(36.3)	(-0.1)	(113.5)	(86.8)
$V_n(5)$	0.92	1.07	1.83	1.45
(%–Bias)	(36.1)	(-6.4)	(130.7)	(95.7)
$V_n(6)$	0.92	1.10	1.72	1.40
(%–Bias)	(35.3)	(9.3)	(145.7)	(102.7)
$V_n(7)$	0.93	1.12	1.63	1.36
(%–Bias)	(34.4)	(-10.8)	(159.0)	(107.9)
$V_n(8)$	0.94	1.14	1.56	1.34
(%–Bias)	(33.2)	(-12.7)	(170.9)	(111.5)

[1] Under the null hypothesis H (conditions [A1], [A2'], and [A3]-[A5]), the limiting distribution of $V_n(q)$ is the range of a Brownian bridge, which has a mean of $\sqrt{\pi/2}$. Fractiles are given in table 1a; the 95 percent confidence interval with equal probabilities in both tails is (0.809, 1.862). Entries in the %-Bias rows are computed as $\{[\tilde{V}_n/V_n(q)]^{1/2} - 1\} \cdot 100$ and are estimates of the bias of the classical R/S statistic in the presence of short-term dependence. Asterisks indicate significance at the 5 percent level.

Source: Authors

Table 3

Finite sample distribution of the modified R/S statistic under i.i.d. and ARMA(2,2) null hypotheses for the first difference of real log-GNP. The Monte Carlo experiments under the two null hypotheses are independent and consist of 10,000 replications each. Parameters of the i.i.d. simulations were chosen to match the sample mean and variance of quarterly real GNP growth rates from 1947:IQ to 1987:IVQ; parameters of the ARMA(2,2) were chosen to match point estimates of an ARMA(2,2) model fitted to the same data set. Entries in the column labeled "q" indicate the number of lags used to compute the R/S statistic. A lag of zero corresponds to Mandelbrot's classical R/S statistic, and a non-integer lag value corresponds to the average (across replications) lag value used according to Andrews' (1991) optimal lag formula. Standard errors for the empirical size may be computed using the usual normal approximation; they are 9.95×10^{-4}, 2.18×10^{-3}, and 3.00×10^{-3} for the 1, 5, and 10 percent tests, respectively.

i.i.d. Null Hypothesis:

n	q	Min	Max	Mean	S. D.	Size 1%-Test	Size 5%-Test	Size 10%-Test
163	0	0.522	2.457	1.167	0.264	0.022	0.081	0.138
163	1.5	0.525	2.457	1.171	0.253	0.015	0.069	0.121
163	1	0.533	2.423	1.170	0.254	0.016	0.069	0.125
163	2	0.564	2.326	1.174	0.246	0.011	0.058	0.111
163	3	0.602	2.221	1.179	0.239	0.009	0.046	0.097
163	4	0.641	2.136	1.184	0.232	0.006	0.036	0.082
163	5	0.645	2.087	1.189	0.225	0.004	0.030	0.071
163	6	0.636	2.039	1.193	0.219	0.002	0.024	0.061
163	7	0.648	1.989	1.198	0.213	0.000	0.018	0.050
163	8	0.657	1.960	1.203	0.207	0.000	0.015	0.040

ARMA(2,2) Null Hypothesis:

n	q	Min	Max	Mean	S. D.	Size 1%-Test	Size 5%-Test	Size 10%-Test
163	0	0.746	3.649	1.730	0.396	0.175	0.340	0.442
163	6.8	0.610	2.200	1.177	0.229	0.009	0.041	0.084
163	1	0.626	3.027	1.439	0.321	0.034	0.110	0.182
163	2	0.564	2.625	1.273	0.279	0.010	0.054	0.111
163	3	0.550	2.412	1.202	0.257	0.012	0.055	0.106
163	4	0.569	2.294	1.180	0.244	0.012	0.054	0.102
163	5	0.609	2.241	1.178	0.236	0.010	0.048	0.093
163	6	0.616	2.181	1.180	0.229	0.008	0.040	0.082
163	7	0.629	2.109	1.180	0.222	0.006	0.034	0.073
163	8	0.644	2.035	1.179	0.215	0.005	0.030	0.066

Source: Authors

Table 4

Power of the modified R/S statistic under a Gaussian fractionally differenced alternative with differencing parameters d = 1/3, -1/3. The Monte Carlo experiments under the two alternative hypotheses are independent and consist of 10,000 replications each. Parameters of the simulations were chosen to match the sample mean and variance of quarterly real GNP growth rates from 1947:IQ to 1987:IVQ. Entries in the column labeled "q" indicate the number of lags used to compute the R/S statistic; a lag of zero corresponds to Mandelbrot's classical R/S statistic, and a non-integer lag value corresponds to the average (across replications) lag value used according to Andrews' (1991) optimal lag formula.

d = 1/3:

n	q	Min	Max	Mean	S. D.	Power 1%-Test	Power 5%-Test	Power 10%-Test
163	0	0.824	4.659	2.370	0.612	0.637	0.778	0.839
163	6.0	0.702	2.513	1.524	0.286	0.017	0.126	0.240
163	1	0.751	3.657	2.004	0.478	0.416	0.587	0.680
163	2	0.721	3.140	1.811	0.409	0.254	0.448	0.545
163	3	0.708	2.820	1.688	0.363	0.141	0.331	0.440
163	4	0.696	2.589	1.600	0.330	0.068	0.234	0.350
163	5	0.700	2.417	1.534	0.304	0.027	0.158	0.271
163	6	0.700	2.297	1.482	0.282	0.008	0.096	0.201
163	7	0.699	2.195	1.440	0.264	0.001	0.056	0.141
163	8	0.694	2.107	1.405	0.249	0.000	0.027	0.097

d = -1/3:

n	q	Min	Max	Mean	S. D.	Power 1%-Test	Power 5%-Test	Power 10%-Test
163	0	0.352	1.080	0.614	0.103	0.849	0.956	0.981
163	4.1	0.449	1.626	0.838	0.142	0.211	0.456	0.600
163	1	0.416	1.251	0.708	0.116	0.587	0.811	0.895
163	2	0.456	1.344	0.779	0.125	0.350	0.631	0.758
163	3	0.512	1.467	0.837	0.132	0.194	0.458	0.612
163	4	0.546	1.545	0.887	0.137	0.100	0.309	0.471
163	5	0.564	1.667	0.931	0.141	0.046	0.200	0.334
163	6	0.600	1.664	0.970	0.144	0.019	0.124	0.236
163	7	0.638	1.731	1.007	0.147	0.008	0.074	0.158
163	8	0.652	1.775	1.041	0.149	0.004	0.041	0.103

Source: Authors