# LONG MEMORY AT THE LONG RUN AND AT THE CYCLICAL FREQUENCIES: MODELLING REAL WAGES IN ENGLAND, 1260 -1994

**Guglielmo Maria Caporale**

**Brunel University, London**

**Luis A. Gil-Alana**[*]

**University of Navarra**

**October 2004**

## Abstract

*This paper examines historical data on daily real wages in England for the time period 1260-1994 by means of new statistical techniques suitable for modelling long memory both at the long run and the cyclical frequencies. Specifically, it uses a procedure due to Robinson (1994) which is based, for the cyclical component, on Gegenbauer processes. We test for the presence of unit (and fractional) roots at both the zero and the cyclical frequencies, and find that the root at the zero frequency plays a much more important role than the cyclical one, though the latter frequency also has a component of long memory behaviour. It also appears that the trending (zero frequency) component is nonstationary while the cyclical one is stationary, with shocks having permanent effects on the former, but transitory effects on the latter. Similar conclusions are reached when allowing for a break in 1875 (the beginning of the Second Industrial Revolution).*

***Keywords:*** *Gegenbauer Processes, Fractional Cycles, Long Memory*

***JEL classification:*** *C22*

***Corresponding author:*** *Professor Guglielmo Maria Caporale, Brunel Business School, Brunel University, Uxbridge, Middlesex UB8 3PH, UK.*
*Email: Guglielmo-Maria.Caporale@brunel.ac.uk*

[*] We are grateful to Bernd Fitzenberger and two anonymous referees for useful comments and suggestions. The second named author also gratefully acknowledges financial support from the Minsterio de Ciencia y Tecnologia (SEC2002-01839, Spain).

## 1. Introduction

The importance of adequately modelling the cyclical component of macroeconomic time series has been widely documented in the literature. Following the seminal study of Burns and Mitchell (1946), numerous authors have investigated empirically cyclical fluctuations, analysing their varying degree of smoothness (see, e.g., Romer, 1986, 1994, and Diebold and Rudebusch, 1992), their asymmetry (see, inter alia, Neftci, 1984, and Hamilton, 1989), or their length (see Baxter and King, 1999, Canova, 1998, and King and Rebelo, 1999). They typically find that business cycles after the Second World War have lower amplitude and longer duration, that recessions are deeper and shorter than expansions, and that the average length of a cycle is six years. More recent research has also shown that simple linear ARIMA(1,1,0) models replicate business cycles at least as accurately as more complex linear or nonlinear specifications (see Hess and Iwata, 1997), and has proposed the use of fractionally ARIMA (ARFIMA) models to characterise cyclical fluctuations (see Candelon and Gil-Alana, 2004). Furthermore, it has been argued that cycles should be modelled as an additional component to the trend and the seasonal structure of the series (see Harvey, 1985 and Gray *et al*, 1989). All these studies deal with US data, and hence their conclusions on duration, dependence and the lengths of the phase of the cycle are not necessarily applicable to other countries.

This paper uses UK data and focuses on real wages, whose cyclical behaviour has often been examined by macroeconomists (see, e.g., Blanchard and Katz, 1999). A common finding is that real wages are procyclical rather than countercyclical, which is not consistent with the Keynesian characterisation of the transmission mechanism of demand shocks to output. Neither can their observed fluctuations be easily reconciled with competitive equilibrium business cycles, with technological shocks driving the cycle (see Blanchard and Fischer, 1989). In the present study, we analyse historical data on real daily wages in England for the time period 1260-1994. Clearly, over such a long time period, structural breaks are likely to have occurred between different "regimes" in wage determination (e.g., before and after the creation of trade unions, or different periods in the unionisation movement), and the empirical analysis should allow for them (as we do at the end of Section 3). Further, given the fact that fundamental changes have taken place in the process of wage determination, it would be inappropriate to refer to the same theoretical model to interpret the findings over the whole sample. In more recent period, the theoretical framework generally used in the UK is a competing-claims model of a unionised economy with imperfect competition, with wages being determined through collective bargaining and prices being set by imperfectly competitive firms (see Layard, Nickell and Jackman, 1994). Consequently, wages and prices are modelled as a simultaneous dynamic system, which includes cointegrating relationships. This requires that all series must be I(1), which is the conclusion generally reached in applied studies (see, e.g., Bardsen and Fisher, 1999, and the references therein).[1]

However, the tests normally carried out do not allow for the possibility of fractional orders of integration. By contrast, the present study tests for the presence of both fractional and integer orders of integration. Furthermore, in view of the earlier literature on business cycles, it adopts a modelling approach which, instead of considering exclusively the component affecting the long-run or zero frequency, also takes into account the cyclical structure, which is modelled as a Gegenbauer process.

The structure of the paper is as follows. Section 2 briefly describes the Robinson (1994) tests used for the empirical analysis. Section 3 discusses an application to annual data on daily wages in England for the time period 1260 –1994, while Section 4 contains some concluding comments.

[1] An I(1) process is defined as a process that requires first differences to achieve I(0). A proper definition of an I(0)

## 2. The testing procedure

Robinson (1994) considers the regression model:

$$y_t \;=\; \beta' z_t \;+\; x_t \qquad t \;=\; 1, 2, \ldots, \tag{1}$$

where $y_t$ is a given raw time series; $z_t$ is a (kx1) vector of deterministic regressors that may include, for example, an intercept, (e.g., $z_t \equiv 1$), or an intercept and a linear time trend, (in case of $z_t = (1,t)'$); $\beta$ is a (kx1) vector of unknown parameters; and the regression errors $x_t$ are such that:

$$\rho(L;\,\theta)\; x_t \;=\; u_t \qquad t \;=\; 1, 2, \ldots, \tag{2}$$

where $\rho$ is a given function which depends on L (the lag-operator, $Lx_t = x_{t-1}$), and the parameter vector $\theta$, taking the form:

$$\rho(L;\,\theta) \;=\; (1-L)^{d_1+\theta_1}\,(1-L^4)^{d_s+\theta_s}\,\prod_{j=2}^{p}(1-2\cos w\,L+L^2)^{d_j+\theta_j}\;, \tag{3}$$

for real given numbers $d_1$, $d_s$, $d_2$, … $d_p$, integer p, and where $(u_t)_t$ is an I(0) process, defined for the purpose of the present paper as a covariance stationary process with spectral density function that is positive and finite at any frequency on the spectrum.[2] Under the null hypothesis:

$$H_o: \;\; \theta \;=\; 0 \tag{4}$$

equation (3) becomes:

$$\rho(L;\,\theta=0) \;=\; \rho(L) \;=\; (1-L)^{d_1}\,(1-L^4)^{d_s}\,\prod_{j=2}^{p}(1-2\cos w\,L+L^2)^{d_j}\;. \tag{5}$$

This is a very general specification that makes it possible to consider different models under the null. For example, if $d_1 = 1$ and $d_s = d_j = 0$ for $j \geq 2$, we have the classical unit-root model (Dickey and Fuller, 1979, Phillips and Perron, 1988, etc.), and, if $d_1$ is a real value, the fractional models examined in Diebold and Rudebusch (1989), Baillie (1996) and others. Similarly, if $d_s$ = 1 and $d_j = 0$ for $j \geq 1$, we have the seasonal unit-root model (Dickey, Hasza and Fuller, 1984, Hyllerberg *et al.*, 1990, etc.) and, if $d_s$ is real, the seasonal fractional model analysed in Porter-

---

process is given in Section 2.

Hudak (1990). Finally, if $d_1 = d_s = 0$, we have the 1- and k-factor Gegenbauer processes studied in Gray *et al.* (1989, 1994) and in Ferrara and Guegan (2001) respectively.

In this paper we are concerned with both the long run and the cyclical structure of the series, and thus we assume that $d_s = 0$ and $p = 2$. In such a case, equation (3) can be written as:

$$\rho(L;\theta) = (1-L)^{d_1+\theta_1}(1-2\cos wL+L^2)^{d_2+\theta_2} \qquad (6)$$

and, similarly, equation (5) becomes:

$$\rho(L) = (1-L)^{d_1}(1-2\cos wL+L^2)^{d_2}. \qquad (7)$$

Here, $d_1$ represents the degree of integration at the zero frequency (i.e., the stochastic trend), while $d_2$ affects the cyclical component of the series.

Starting with the trend, the first polynomial in equation (7) can be expressed in terms of its Binomial expansion, such that:

$$(1-L)^{d_1} = \sum_{j=0}^{\infty}\binom{d_1}{j}(-1)^j L^j = 1 - d_1 L + \frac{d_1(d_1-1)}{2}L^2 - \dots,$$

for all real $d_1$. If $d_1 \in (0, 0.5)$, the series is covariance stationary (in relation to the zero frequency), whilst if $d_1 \in [0.5, 1)$ it is no longer stationary but still mean reverting, with the effects of the shocks dying away in the long run.

Similarly, for the cyclical component Gray *et al.* (1989, 1994) showed that the second polynomial in equation (7) can be expressed in terms of the Gegenbauer polynomial $C_{j,d_2}(\mu)$ such that, denoting $\mu = \cos w$, for all $d_2 \neq 0$,

$$(1-2\mu L+L^2)^{-d_2} = \sum_{j=0}^{\infty} C_{j,d_2}(\mu)L^j, \qquad (8)$$

where

$$C_{j,d_2}(\mu) = \sum_{k=0}^{[j/2]}\frac{(-1)^k (d_2)_{j-k}(2\mu)^{j-2k}}{k!(j-2k)!}; \qquad (d_2)_j = \frac{\Gamma(d+j)}{\Gamma(d)},$$

[2] In other words, $0 < f(\lambda) < \infty$, where $f(\lambda)$ is the spectral density function of the process.

where Γ(x) represents the Gamma function. They also showed that the series is then stationary (with respect to the cyclical part) if $|\mu| < 1$ and $d_2 < 0.50$ or if $|\mu| = 1$ and $d_2 < 0.25$.

We next describe the test statistic. We observe $\{(y_t, z_t), t = 1,2,\ldots n\}$, and assume that the I(0) process $(u_t)_t$ in equation (2) has spectral density given by:

$$f(\lambda;\tau) = \frac{\sigma^2}{2\pi} \, g(\lambda;\tau), \qquad -\pi < \lambda \leq \pi ,$$

where the scalar $\sigma^2$ is known and g is a function of known form, which depends on frequency λ and the unknown (qx1) vector τ that controls the short run components of the series. Under $H_o$ (4), the residuals in the equations (1), (2) and (6) are:

$$\hat{u}_t = (1 - L)^{d_1}(1 - 2\cos wL + L^2)^{d_2} \, y_t - \hat{\beta}' s_t , \tag{9}$$

$$\hat{\beta} = \left( \sum_{t=1}^{n} s_t \, s_t' \right)^{-1} \sum_{t=1}^{n} s_t (1-L)^{d_1} (1 - 2\cos wL + L^2)^{d_2} \, y_t , \qquad s_t = (1-L)^{d_1}(1 - 2\cos wL + L^2)^{d_2} \, z_t .$$

Unless g is a completely known function (e.g., g ≡ 1, as when $u_t$ is white noise), we need to estimate the nuisance parameter τ, for example by $\hat{\tau} = \arg min_{\tau \in T^*} \, \sigma^2(\tau)$, where $T^*$ is a suitable subset of $R^q$ Euclidean space, and

$$\sigma^2(\tau) = \frac{2\pi}{n} \sum_{s=1}^{n-1} g(\lambda_s;\tau)^{-1} I_{\hat{u}}(\lambda_s), \ \text{with} \ \ I_{\hat{u}}(\lambda_s) = \left| (2\pi n)^{-1/2} \sum_{t=1}^{n} \hat{u}_t \, e^{i\lambda_s t} \right|^2 ; \qquad \lambda_s = \frac{2\pi s}{n},$$

where s = 1, 2, …, and n is the sample size. The test statistic, which is derived through the Lagrange Multiplier (LM) principle takes the form:

$$\hat{R} = \frac{n}{\hat{\sigma}^4} \hat{a}' \hat{A}^{-1} \hat{a}, \tag{10}$$

$$\hat{a} = \frac{-2\pi}{n} \sum_{s}^{*} \psi(\lambda_s) \, g(\lambda_s;\hat{\tau})^{-1} I(\lambda_s); \qquad \hat{\sigma}^2 = \sigma^2(\hat{\tau}) = \frac{2\pi}{n} \sum_{s=1}^{n-1} g(\lambda_s;\hat{\tau})^{-1} I(\lambda_s),$$

$$\hat{A} = \frac{2}{n} \left( \sum_{s}^{*} \psi(\lambda_s)\psi(\lambda_s)' - \sum_{s}^{*} \psi(\lambda_s)\hat{\varepsilon}(\lambda_s)' \left( \sum_{s}^{*} \hat{\varepsilon}(\lambda_s)\hat{\varepsilon}(\lambda_s)' \right)^{-1} \sum_{s}^{*} \hat{\varepsilon}(\lambda_s)\psi(\lambda_s)' \right)$$

$$\psi(\lambda_s)' = [\psi_1(\lambda_s), \psi_2(\lambda_s)]; \quad \hat{\varepsilon}(\lambda_s) = \frac{\partial}{\partial \tau} \log g(\lambda_s; \hat{\tau});$$

$$\psi_1(\lambda_s) = \log\left|2 \sin\frac{\lambda_s}{2}\right|; \qquad \psi_2(\lambda_s) = \log\left|2\left(\cos\lambda_s - \cos w\right)\right|,$$

and the summation on * in the above expressions are over $\lambda \in M$ where $M = \{\lambda: -\pi < \lambda < \pi, \lambda \notin (\rho_l - \lambda_f, \rho_l + \lambda_f), f = 1, 2, \ldots, m\}$, such that $\rho_f$, $f = 1, 2, \ldots, m < \infty$ are the distinct poles of $\psi(\lambda)$ on $(-\pi, \pi]$.

Under $H_o$ (4), Robinson (1994) established that, under certain regularity conditions:

$$\hat{R} \to_d \chi_2^2, \qquad as \quad n \to \infty, \tag{11}$$

where "$\to_d$" means convergence in distribution. These conditions are very mild and concern technical assumptions which are satisfied by $\psi_1(\lambda)$ and $\psi_2(\lambda)$. These assumptions are required to justify approximating integrals by sums when deriving the test statistic, and are satisfied not only by the model in equations (2) and (6) but also by virtually any other model an applied researcher could think of (see Robinson, 1994, p. 1424). Thus, unlike in other procedures, we are in a classical large-sample testing situation, and furthermore the tests are efficient in the Pitman sense against local departures from the null.[3] It is important to note that we test $H_o$ (4) for any real values $d_1$ and $d_2$, and thus the procedure is asymptotically valid in both stationary and nonstationary contexts. In fact, this is one of the distinguishing features of Robinson's (1994) tests compared with other methods, where preliminary integer differentiation is required to achieve stationarity and invertibility.

## 3. An empirical application to real daily wages in England, 1260 –1994

The time series we consider is real daily wages in England, annually, from 1260 to 1994, which can be obtained from Madridakis *et al* (1998), and is also downloadable from the website:

*http://www-personal.buseco.monash.edu.au/~hyndman/TSDL/data/wagesuk.dat*

**(Insert Figure 1 about here)**

Figure 1 displays the data for the original series. We can see that the values increase over time, especially in the twentieth century. The behaviour of this series at the long-run or zero frequency has been recently examined in Gil-Alana (2004). He uses another version of the tests of Robinson (1994), testing $H_o$ (4) in a model given by the equations (1) and (2), with $z_t = (1,t)'$, $t \geq 1$, $(0,0)'$ otherwise, and $\rho(L; \theta) = (1 - L)^{d+\theta}$. He tests $H_o$ for values d = 0, (0.01), 2, and different types of disturbances. His results can be summarised as follows: if $u_t$ is white noise, the unit root is excluded from the intervals, in favour of higher orders of integration, with d oscillating between 1.02 and 1.13. However, if the disturbances are autocorrelated (either with AR or Bloomfield $u_t$),[4] the unit root cannot be rejected, with d ranging now between 0.88 (Bloomfield (m = 1) $u_t$ with an intercept and/or a time trend) and 1.11 (no regressors and AR(2) $u_t$). A similar result is obtained in the same paper when using a semiparametric (Robinson, 1995) procedure, finding strong evidence of unit roots.

The above two approaches to investigating the long-run behaviour of a time series consist of testing a parametric model for the series and estimating a semiparametric one, relying on the long run-implications of the estimated model. The primary advantage is the precision gained by providing all the information about the series through the parameter estimates. A drawback is that these estimates are sensitive to the class of models considered, and may be misleading because of misspecification. It is well known that the issue of misspecification can never be settled conclusively in the case of parametric (or semiparametric) models. However, the problem can be addressed by considering a larger class of models.

---

[3] In other words, if the tests are implemented against local departures of the form: $H_a$: $\theta = \delta n^{-1/2}$, for $\delta \neq 0$, the limit distribution is a $\chi_2^2(v)$ with a non-centrality parameter v, which is optimal under Gaussianity of $u_t$.

[4] The Bloomfield's (1973) model is a non-parametric approach to modelling the I(0) disturbances, which approximates the ARMA structure and fits fairly well into the present version of the tests.

For this purpose, let us consider now the model given by equations (1) and (2), with ρ(L; θ) as in equation (6). Thus, under $H_o$ (4), the model becomes:

$$y_t \;=\; \beta_0 \;+\; \beta_1 t \;+\; x_t, \qquad t \;=\; 1, 2, ... \qquad (12)$$

$$(1 - L)^{d_1}\; (1 - 2\cos w\, L + L^2)^{d_2}\; x_t \;=\; u_t, \qquad t \;=\; 1, 2, ...., \qquad (13)$$

and if $d_2 = 0$, the model reduces to the case previously mentioned of long memory exclusively at the long-run or zero frequency. We assume that $w = w_r = 2\pi/r$, r indicating the number of time periods per cycle.[5]

**(Insert Table 1 about here)**

We computed the statistic $\hat{R}$ given by equation (10) for values of $d_1$ and $d_2$ = 0.10, (0.10), 2, and r = 1, …, n/2, assuming that $u_t$ is white noise. However, instead of presenting the results for all values of r, $d_1$ and $d_2$, we only report in Table 1 cases of non-rejection values at the 95% significance level. We can see that the values of r where $H_o$ cannot be rejected are between 5 and 14 years, with $d_1$ equal to 1 and 1.1, and $d_2$ = 0.10 and 0.20. Thus, we observe higher orders of integration at the long-run or zero frequency compared with the cyclical one, though the latter frequency also seems to have a component with a long memory behaviour. Also, the fact that $H_o$ (4) is decisively rejected for values of r smaller than 5 and higher than 14 and for most of the values of $d_1$ and $d_2$ suggests that the optimal power properties of Robinson's (1994) tests are supported by their reasonable performance against non-local alternatives and different values of r.

In order to be even more precise about the non-rejection values of $d_1$ and $d_2$, we re-computed the tests but this time for a shorter grid, with $d_1$, $d_2$ = 0.01, (0.01), 2. Figure 2 displays the regions of ($d_1$, $d_2$) values where $H_o$ (4) cannot be rejected at the 95% level, for the case of an intercept, with r = 6. Other values of r were also employed and very similar results were

[5] The series analysed in the paper is a logarithmic transformation of real daily wages, though similar results were obtained when using the original data.

obtained. We can see (in the left hand side of the figure) that, if $(u_t)_t$ is a white noise process, $d_1$ ranges between 0.95 and 1.15, and $d_2$ is between 0.05 and 0.20. This indicates that the series is nonstationary with respect to the long-run or zero frequency and stationary with respect to the cyclical structure.

We also performed the tests allowing for autocorrelated disturbances, with $(u_t)_t$ which follows an AR(1) process. Higher AR orders were also employed and the results did not substantially differ from those reported. In general, we observe a higher proportion of non-rejection values compared with the case of white noise $(u_t)_t$, though the same qualitative conclusions are reached, with $d_1$ oscillating between 0.9 and 1.1, and $d_2$ ranging between 0 and 0.2. Note that in this case the null hypothesis cannot be rejected for some values of $d_1$ with $d_2 = 0$. However, it should also be noted that these hypotheses are "less clearly non-rejected"[6] than when $d_2$ is positive, suggesting that a component of long memory is present at the cyclical frequency. In view of these findings, it seems clear that the root at the zero frequency plays a much more important role than the one at the cyclical frequency, though the latter also appears to have a component of long memory (stationary) behaviour.

**(Insert Figures 2 and 3 about here)**

The final issue to investigate is the possibility of a structural break in the data. In particular, we consider the case of a break occurring at the beginning of the Second Industrial Revolution (SIR, 1875). We chose that date because it corresponds to the period witnessing the move from "New Model" unionism, which had characterised the years from the late 1840s until then (with an increased degree of nationalisation and centralisation in the unions representing workers in the main profitable industries), to "New" unionism (which extended to formerly unorganised workers, and led to a mass labour movement). Also, new important legislation was

[6] By "less clearly non-rejected" we mean that the value of the test statistic is closer to the rejection critical values.

introduced around that time (e.g. the Employers and Workmen Act in 1875 gave workers the right to sue their employers for breach of contract).[7]

Robinson's (1994) procedure allows us to include dummy variables for the break in the regression model given by equation (1). Thus, we can consider the model

$$y_t \;=\; \beta_0 \;+\; \beta_1 D_t \;+\; x_t, \qquad t \;=\; 1,\, 2,\, \ldots$$

and (13), where $D_t = 1$ for $t < T_b$, and $(t - T_b)$ for $t \geq T_b$, that is, we assume a slope dummy at $T_b$ = 1875. Figure 2 displays the results for the two cases of white noise and AR(1) disturbances. We can see that similar conclusions are obtained with the values of $d_1$ ranging between 0.95 and 1.15, and $d_2$ from 0 to 0.22. We also considered other types of breaks (e.g. level shift dummies) and other breakpoints (e.g. First Industrial Revolution, World War II), but the corresponding coefficients were not found to be significantly different from zero.

## 4. Conclusions

In this paper we have examined the time series behaviour of real wages in England for the time period 1260 - 1994 by means of new statistical techniques based on long memory processes. Specifically, we have used a procedure due to Robinson (1994) that has enabled us to test for roots not only at zero but also at the cyclical frequencies. These tests have standard null and local limit distributions and can easily be applied to raw time series. Another version of Robinson's (1994) tests which is appropriate for investigating the root at the long-run or zero frequency was implemented in Gil-Alana (2004). His results show that d is close to 1, implying that a unit root is present. Similar results were obtained in the same paper when using a semiparametric procedure (QMLE – see Robinson, 1995).

However, the non-rejected values obtained at the zero frequency could partly be due to the fact that no attention was paid to other possible (cyclical) frequencies of the process. In fact,

[7] See TUC History online, http://www.unionhistory.info/timeline/timeline.php, for further details.

visual inspection of the periodogram of the first differenced data (not reported) shows a peak at a frequency close to $2\pi/6$, suggesting that a component of long memory behaviour might still be present in the cyclical component. Thus, we adopted a method suitable for simultaneously testing for the presence of roots at the zero and cyclical frequencies. For the latter frequency, the model is based on Gegenbauer processes. The results suggest that the periodicity of the series ranges between 5 and 14 years. This finding is in line with those of most of the empirical literature on cycles reporting a periodicity of about six years (see, e.g., Baxter and King, 1999, Canova, 1998, and King and Rebelo, 1999). However, it is hardly possible to compare the growth cycle in post-war US GNP analysed in those papers, with the cycle of real wages in England from 1260 examined in the present paper. An appropriate comparison would be that between cycles in output and wages in the same country and over the same time period, but this is beyond the scope of our study, which focuses on the behaviour of wages only. Their order of integration seems to fluctuate around 1 at the long-run frequency and between 0 and 0.2 at the cyclical one, implying that the series is nonstationary with respect to the zero frequency but stationary with respect to the cyclical frequency. Similar conclusions were obtained when a break at the beginning of the Second Industrial Revolution (1875) was introduced in the model.

As for the policy implications of our findings, and concentrating on the more recent period of a unionised economy, it is generally argued that supply side policies, including those towards wage determination (such as the reform of wage bargaining, or subduing earnings growth) are most effective in reducing unemployment (see Layard, Nickell and Jackman, 1994), whilst demand management policies (such as a looser fiscal stance or a lower exchange rate) have prolonged, but not permanent effects (see Barrell, Caporale and Sefton, 1994). However, our analysis suggests that the time series properties of wages are more complex than previously found, as they are characterised by both a long-run, nonstationary component (permanently affected by shocks) and by a cyclical, stationary one (with shocks having only temporary effects). Such complex behaviour should be taken into account when devising policies to reduce wage inflation and to combat unemployment.

## References

Baillie, R.T., 1996, Long memory processes and fractional integration in econometrics, Journal of Econometrics 73, 5-59.

Bardsen, G. and P.G. Fisher, 1999, Economic theory and econometric dynamics in modelling wages and prices in the United Kingdom, Empirical Economics 24, 3, 483-507.

Barrell, R., Caporale, G.M., and J. Sefton, 1994, Prospects for European unemployment, in J. Michie and J. Grieve Smith (eds.), Unemployment in Europe, 32-44, Academic Press, Cambridge.

Baxter, M. and R.G. King, 1999, Measuring business cycles approximate band-pass filters for economic time series, The Review of Economics and Statistics 81, 575-593.

Blanchard, O.J. and S. Fischer, 1989, Lectures on Macroeconomics, MIT Press, Cambridge, Massachusetts.

Blanchard, O.J. and L. Katz, 1999, Wage dynamics: reconciling theory and evidence, American Economic Review 89, 2, 69-74.

Bloomfield, P., 1973, An exponential model in the spectrum of a scalar time series, Biometrika 60, 217-226.

Burns, A. and W.C. Mitchell, 1946, Measuring Business Cycles, NBER, New York.

Canova, F., 1998, Detrending and business cycle facts. A user's guide, Journal of Monetary Economics 41, 533-540.

Candelon, B. and L.A. Gil-Alana, 2004, Fractional integration and business cycle features, Empirical Economics 29, 343-359.

Dickey, D. A. and W. A. Fuller, 1979, Distribution of the estimators for autoregressive time series with a unit root, Journal of the American Statistical Association 74, 427-431.

Dickey, D. A., D. P. Hasza and W. A. Fuller, 1984, Testing for unit roots in seasonal time series, Journal of the American Statistical Association 79, 355-367.

Diebold, F.X. and G.D. Rudebusch, 1992, Have post-war economic fluctuations been stabilised?, The American Economic Review 82, 993-1005.

Ferrara, L. and D. Guegan, 2001, Forecasting with k-factor Gegenbauer processes: Theory and applications, Journal of Forecasting 20, 581-601.

Gil-Alana, L.A. 2004, A re-examination of historical real daily wages in England, forthcoming in Journal of Policy Modelling.

Gray, H.L., Yhang, N. and Woodward, W.A., 1989, On generalized fractional processes, Journal of Time Series Analysis 10, 233-257.

Gray, H.L., Yhang, N. and Woodward, W.A., 1994, On generalized fractional processes. A correction, Journal of Time Series Analysis 15, 561-562.

Hamilton, J.D., 1989, A new approach to the economic analysis of nonstationary time series and the business cycle, Econometrica 57, 357-384.

Harvey, A., 1985, Trends and cycles in macroeconomic time series, Journal of Business and Economics and Statistics 3, 216-227.

Hess, G.D. and S. Iwata 1997, Asymmetric persistence in GDP? A Deeper look at depth, Journal of Monetary Economics 40, 535-554.

Hyllerberg, S., R.F. Engle, C.W.J. Granger and B.S. Yoo, 1990, Seasonal integration and cointegration, Journal of Econometrics 44, 215-238.

King, R.G. and S.T. Rebelo, 1999, Resucitating real business cycles, in J.B. Taylor and M. Woodford eds., Handobook in Econometrics, Vol. 1, 928-1001.

Layard R., S. Nickell and R. Jackman, 1994, The Unemployment Crisis, Oxford University Press, Oxford.

Madridakis, S., Wheelwright, S.C. and R. Hyndman, 1998, Forecasting Methods and Applications, 3rd Edition, John Wiley.

Neftci, S. 1984, Are economic time series asymmetric over the business cycle?, Journal of Political Economy 94, 1096-1109.

Phillips, P.C.B. and P. Perron, 1988, Testing for a unit root in a time series regression, Biometrika 75, 335-346.

Porter-Hudak, S., 1990, An application of the seasonally fractionally integrated model to the monetary aggregates, Journal of the American Statistical Association 85, 338-344.

Robinson, P.M., 1994, Efficient tests of nonstationary hypotheses, Journal of the American Statistical Association 89, 1420-1437.

Robinson, P.M., 1995, Log-periodogram regression of time series with long range dependence, Annals of Statistics 23, 1048-1072.

Romer, C. 1986, Spurious volatility in historical unemployment data, Journal of Political Economy 94, 1-36.

Romer, C., 1994, Remeasuring business cycles, Journal of Economic History 54, 573-609.

**FIGURE 1**

**Historical real daily wages in England (1260-1994)**

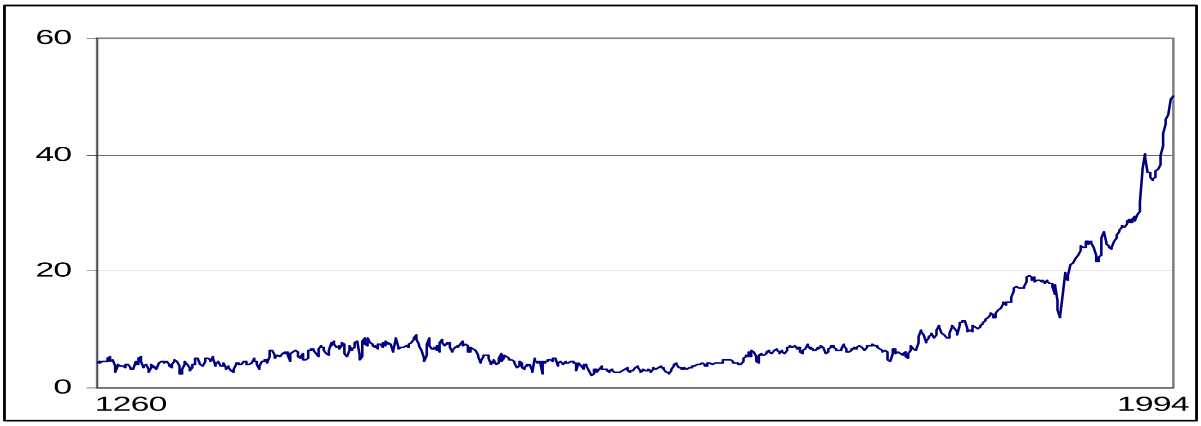



**TABLE 1**

Testing $H_o$ (4) in (12) and (13) with white noise $u_t$

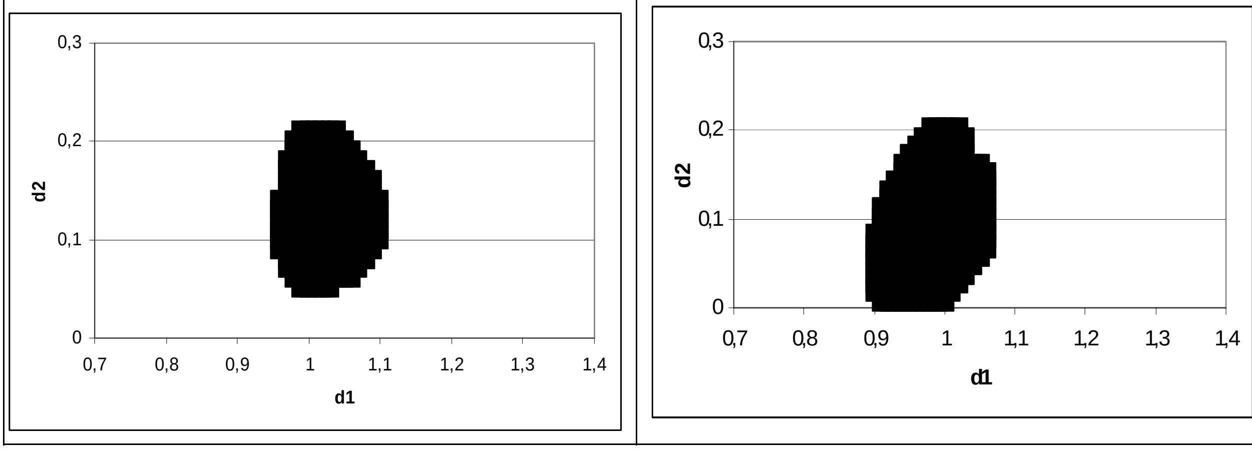
| r | $d_1$ | $d_2$ | No regressors | An intercept | A linear trend |
|---|---|---|---|---|---|
| 5 | 1.00 | 0.10 | 9.134 | **2.776** | **2.817** |
| 5 | 1.10 | 0.10 | **1.371** | 8.098 | **3.503** |
| 5 | 1.10 | 0.20 | **5.619** | 9.003 | 10.119 |
| 6 | 1.00 | 010. | **2.083** | **0.871** | **0.829** |
| 6 | 1.00 | 0.20 | 12.533 | **4.332** | **4.377** |
| 6 | 1.10 | 0.10 | **3.349** | 14.546 | 5.520 |
| 7 | 1.00 | 0.10 | **0.688** | **0.256** | **0.254** |
| 7 | 1.10 | 0.10 | **5.346** | 19.034 | 13.432 |
| 8 | 1.00 | 0.10 | **0.468** | **0.109** | **0.100** |
| 9 | 1.00 | 0.10 | **0.583** | **0.544** | **0.528** |
| 10 | 1.00 | 0.10 | **1.549** | **1.137** | **1.128** |
| 11 | 1.00 | 0.10 | **2.690** | **2.134** | **2.125** |
| 12 | 1.00 | 0.10 | **3.411** | **2.964** | **2.949** |
| 13 | 1.00 | 0.10 | **4.773** | **4.337** | **4.319** |
| 14 | 1.00 | 0.10 | **5.912** | **5.497** | **5.481** |

In bold: Non-rejection values at the 95% significance level.

**FIGURE 2**

Values of $d_1$ and $d_2$ when $H_o$ (4) cannot be rejected at the 95% significance level with r = 6

| White noise disturbances | AR (1) disturbances |
|---|---|

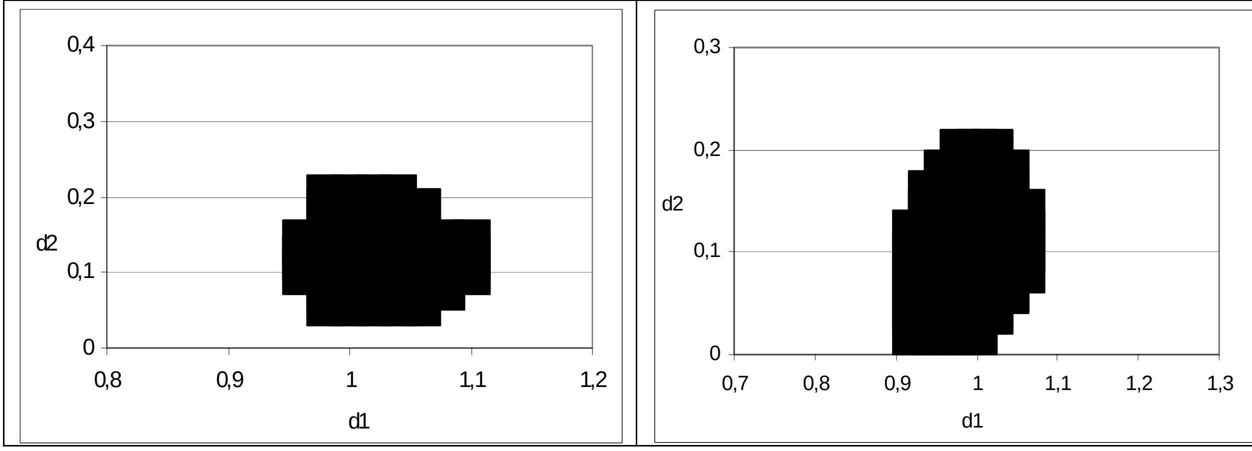



$d_1$ refers to the order of integration at the zero frequency while $d_2$ corresponds to the cyclical component.

**FIGURE 3**

Values of $d_1$ and $d_2$ when $H_o$ (4) cannot be rejected at the 95% significance level with r = 6 and a slope break at 1875

| White noise disturbances | AR (1) disturbances |
|---|---|




$d_1$ refers to the order of integration at the zero frequency while $d_2$ corresponds to the cyclical component.