WHO DID WORSE? A COMPARISON OF US AND BRITISH NON-WHITE UNEMPLOYMENT 1970-1998

by

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Abstract

US and British unemployment rates for non-white males and females are compared over the period 1970-98. Whereas US rates remained fairly steady, there was a marked increase in British non-white unemployment rates. The reasons for this poor performance, relative to the good performance of US non-whites are explored. It is shown that non-white unemployment behaves in different ways across the two countries. For example, British rates rise faster in a recession than white rates, whereas US rates appear not to follow this British hypercyclical pattern.

Key words: non-whites and whites, unemployment, hypercyclicality

JEL Classification: J7

I Introduction

The last 30 years has seen a marked deterioration in the employment prospects of British non-whites both in absolute terms and relative to British whites. By comparison the US record on non-white unemployment is far more impressive. During this period, in Britain the difference between white and non-white unemployment widened considerably, whereas US differentials have remained virtually the same. Britain has had some astonishingly high unemployment rates for certain non-white groups; for example, the 1991 *Population Census* revealed a 50% unemployment rate among young Black African males compared with an overall white rate of 11%. It is surprising that there has been so little social unrest in the face of these numbers.

We are not directly concerned with the underlying determinants of racial differences. Non-white unemployment has already been extensively investigated in Britain by Leslie *et al.* (1998) and Modood (1997) among many others. The US has a similarly extensive literature.[1] The purpose here is to compare the pattern of US non-white unemployment with Britain at an aggregate level.

It has also been suggested that non-white unemployment rates are *hypercyclical* relative to white rates. We may define full hypercyclicality as a situation where non-white unemployment *rises* more rapidly than white unemployment during recessions but *falls* more rapidly during the recovery. Alternatively, non-white

[1] Stratton (1993) is an example of a comparison among US racial groups. Britain's non-whites are mainly (around 70% in 1998) immigrants. Borjas has written extensively on the issue of US immigration of which Borjas (1994) is a useful summary.

employment rates may only be hypercyclical during recessions. We refer to such a situation as partial hypercyclicality. For Britain we find some evidence of partial hypercyclicality but none for the US. For example, according to Modood *et al.* (1997), not only are non-white unemployment rates consistently higher than white rates in Britain but also non-whites suffer disproportionately as unemployment worsens for whites. One obvious question that requires an answer is whether non-whites *benefit* disproportionately as the economic situation improves. That is, is full hypercyclicality present? In Britain this does not appear to be the case. In fact we find evidence for a *ratchet* effect operating in Britain.. That is, although non-whites suffer relative to whites during recession, they are unable to recoup these losses during recovery and, consequently, start a cycle in a worse state than they experienced at the start of the previous cycle.

II The data

The key problem is to construct a reliable series of non-white unemployment for Britain. The US has a regular and consistent series for non-whites dating back to 1954. Unfortunately, British unemployment data for all groups (including whites) have always been somewhat of a muddle. Traditionally the British method was a count of those registered at labor exchanges and after 1982 a count of claimants, that is those actually in receipt of benefits. This gives numbers rather different from the more reliable US survey method. Not all claimants are genuinely unemployed and many genuinely unemployed individuals are not claimants. Furthermore, the British method of recording data has been subject to innumerable changes to the extent that it is not a very reliable labor market indicator. This situation improved in 1973 with the introduction of the *Labor Force Survey* which gave a parallel estimate based on the US survey method. In 1979 the survey also recorded ethnic origin for the first time. Using this and several other sources we have constructed a quarterly series of British white and non-white unemployment for males and females from 1970 - 98. This allows direct comparison with US data for the first time and turns out to be an extremely revealing exercise.[2]

The constructed British series for non-whites are shown alongside US rates in Fig. 1 (males) and Fig. 2 (females) and are not seasonally adjusted. One glaring feature is immediately apparent for both males and females. British rates were below US rates in 1970, but far above in 1998. For example, in the last quarter of 1970 (the start date of our series), British non-white male and female unemployment rates were 4.7% and 6.3 % , while the corresponding US rates were 7.7 % and 9.5%. However, by 1998 the situation had reversed. British rates were 14.1 % and 12.6 %, while US rates were 7.3 % and 7.9%. Overall, average British non-white male unemployment was 3.1% above US rates and female unemployment was 3.0% above US rates over this 29 year comparison period.

[INSERT FIGS 1 - 4 ABOUT HERE]

Of course, another feature of 1970-98 was the fact that British white unemployment rates also deteriorated relative to US white rates - although the average white differences compared with US rates at 2.4% for males and 1.5% for females were somewhat smaller than non-white differences. This might be one reason for the

[2] The method is fully described in Leslie *et al*. (1999). A spreadsheet detailing all the various

deteriorating performance of British non-whites - the fact that the labor market situation was relatively worse for all compared with the US labor market. However, the increasing gap between white and non-white unemployment in Britain suggests something over and beyond the `bad news for everyone' explanation. The non-white/white unemployment differential for British men increased from 1.1% in 1970 to 9.2% in 1998, and from 2.6% to 9.1% for females, once again in strong contrast to the US labor market where these differentials hardly changed at all. Figs 3 and 4 show this ratchet effect.

Another explanation for currently low US rates is a `prison population' effect. The US now locks up a disproportionate number of non-whites, particularly young males. Many of these would have otherwise been recorded as unemployed. In 1998, on average 836 thousand non-white males were unemployed and roughly 700 thousand non-white males were in prison. It takes little imagination to see how this experiment in social engineering could distort the underlying picture. On the other hand, prison stigmatises individuals and many non-whites with a prison record would find it more difficult to find jobs.[3] Also non-white males are disproportionately represented among the British prison population (roughly three to one). In April 1998 122 thousand non-white males were unemployed and 11.3 thousand were in jail, so the in jail/unemployed ratio is a lot smaller.[4]

Eyeballing data provides only *prima facie* evidence for differences and a more systematic analysis is required. The method of attack is to explore the time series

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adjustments is available on request.

[3] Bound and Freeman (1992) argue this as an explanation for high rates among young black males.

properties of the data shown in Figs 1 to 4. If the US and British series differ, then this gives some support to the view that British non-whites fared worse than US non-whites, independently of the fact that British white unemployment was on average greater than in the US.

The approach is non-causal, in that all we can say is that there are significant differences without necessarily explaining why (though we are like everyone else entitled to our personal views on the matter). For example, the US has had an (admittedly 'on-off') tradition of affirmative action which some authorities (Leonard, 1990) claim have reduced non-white unemployment. Britain, despite some strong anti-discrimination legislation, has no such equivalent policy.

III A framework for thought.

It is useful to have a simple model around which to structure the discussion. We adopt the useful distinction developed by Lindbeck and Snower (1988) and split workers into two categories, `insiders' and `outsiders'. Insiders are the privileged group, who through a combination of desirable characteristics, favorable treatment and other intangible advantages are more likely to be employed. Outsiders, by way of contrast, have human capital characteristics that make them less employable, may be discriminated against or are disadvantaged in less tangible ways. For example, in addition to direct discrimination, some outsiders might have a language difficulty or have a taste for isolation which makes them less employable. The important thing is that we need not be terribly precise here other than to imagine two broad categories of

[4] The numbers are estimated from the *Bureau of Justice Statistics Bulletin*, 1999 and *Prison Population*

worker. Non-whites have a larger proportion of outsiders compared with whites and this arises because of a combination of disadvantage and discrimination.

Each group has a natural rate of unemployment around which actual unemployment evolves. This natural rate is given by:

$$U^{*r} = \gamma^{*}\alpha^{r} + \delta^{*}(1 - \alpha^{r}) \qquad r = \text{white, non-white} \qquad (1)$$

where

U^{*r} = overall natural rate of unemployment

γ^{*} = insider natural rate

δ^{*} = outsider natural rate, where $\delta^{*} > \gamma^{*}$.

α^{r} = proportion of workforce who are insiders. It is assumed that this remains constant over the economic cycle. Unemployment varies because the proportion of insiders unemployed at time *t* (γ_t) and outsiders who are unemployed (δ_t) change over the economic cycle. Thus actual unemployment is given by

$$U_t^{r} = \gamma_t\alpha^{r} + \delta_t(1 - \alpha^{r}) \qquad r = \text{white, non-white} \qquad (2)$$

Actual unemployment, relative to the natural rate is given by:

$$\frac{U_t^{r}}{U^{*r}} = r_t = b^{r}[a^{r}\gamma_t + \delta_t] \qquad (3)$$

Brief, Home Office 1999. British figures exclude Scotland.

where $a^r = \dfrac{\alpha^r}{1-\alpha^r}$ *and* $b^r = \dfrac{1}{a^r\gamma^* + \delta^*}$ are time independent parameters, specific to the racial group whereas γ_t and δ_t are time dependent variables independent of the racial group. Non-white unemployment exceeds white unemployment because there is a larger concentration of insiders among the white group compared with non-whites but the same proportion of insiders and outsiders are unemployed across both groups at any point in the economic cycle (a common δ_t and γ_t).

The question here is why non-white unemployment might be hypercyclical relative to whites. The rate of change of r_t is given by

$$\frac{d\ln r_t}{dt} = \frac{a^r \dfrac{d\gamma_t}{dt} + \dfrac{d\delta_t}{dt}}{a^r\gamma_t + \delta_t} \tag{4}$$

Hypercyclicality is concerned with the elasticity of the relative unemployment rates. This is (w = white and n = non-white)

$$\frac{d\ln n_t}{d\ln w_t} = \frac{(\gamma_t a^w + \delta_t)(a^n \dfrac{d\gamma_t}{dt} + \dfrac{d\delta_t}{dt})}{(\gamma_t a^n + \delta_t)(a^w \dfrac{d\gamma_t}{dt} + \dfrac{d\delta_t}{dt})} \tag{5}$$

This elasticity will be greater than one (implying full hypercyclicality) provided that

$$\frac{d\ln\delta_t}{dt} > \frac{d\ln\gamma_t}{dt} \tag{6}$$

In other words if outsiders suffer disproportionately in the downturn and benefit disproportionately in the upturn, unemployment for non-whites will be fully hypercyclical. This seems a plausible hypothesis, but is not inevitable. For example, the US labor market is far more flexible than Britain - real wages have fallen for those in the lowest decile of the earnings distribution unlike the British labor market.[5] So recessions may not impact disproportionately on outsiders in the US, but they may do so in Britain.[6]

Equation (3), with γ_t and δ_t common to whites and non-whites, suggests a simultaneous process between the two groups. Combining groups gives the identity

$$\ln U_t^n \equiv \ln\frac{n_t}{w_t} + \ln\frac{U^{*n}}{U^{*w}} + \ln U_t^w \tag{7}$$

where n_t is non-white unemployment relative to the natural rate and w_t refers to whites relative to the natural rate. The hypercyclical hypothesis can be captured by assuming that $n_t = [w_t]^\beta$, where $\beta > 1$ would reflect hypercyclicality. Thus if $\beta = 1$ the ratio of the unemployment rates would just be the ratio of the respective natural rates. This gives the simple linear form.

$$\ln U_t^n = \ln\frac{U^{*n}}{(U^{*w})^\beta} + \beta\ln U_t^w \tag{8}$$

The hypercyclical parameter β is linked to the insider-outsider analysis by eq.(5), where β will exceed unity if eq.(6) holds and whites have a larger proportion of insiders.[7]

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5 See Juhn *et al.* (1993).

6 The view that the US labor market is more flexible than European labor markets is reviewed in Nickell (1997) and one consequence is a disproportionate impact on racial minorities. High unemployment rates for non-white immigrants is a European phenomenon and is not just confined to Britain. See OECD (1995).

7 OLS may lead to a biased estimate of β in eq.(8). Suppose that U_t^n and U_t^w are determined by $\ln U_t^r = k_1^r + k_2^r X_t + v_t^r$, where r = white, non-white and $k_2^n > k_2^w > 0$ and v_t^r are disturbance terms ~ $(0, \sigma_r^2)$ and $X_t = f(X_{1t}, X_{2t} \ldots)$ is exogenous. Solving yields $\ln U_t^n = \alpha + \beta\ln U_t^w + \varepsilon_t$ where $\alpha = k_1^n - (k_2^n / k_2^w)k_1^w$, $\beta = k_2^n / k_2^w > 1$, and $\varepsilon_t = v_t^n - \beta v_t^w$ Bias occurs if

The discussion so far has assumed full hypercyclicality to be present. That is, while non-white unemployment grows more rapidly in the economic downturn, it falls more rapidly in the recovery period. But there may also occur, independently of hypercyclical effects, a gradual rise (or fall) over time in the differential between non-white and white unemployment rates. We have already alluded to the possibility of ratchet effects in the case of Britain and we try to capture such effects, firstly by the inclusion of linear and quadratic time trends in eq.(8) and, secondly, by the fitting of spline functions.

The spline approach can be seen in the following example. Consider a time period which contains a single turning point for white unemployment, then we can firstly write

$$\ln U_t^n = \alpha_0 + \beta_1 \ln U_t^w + \alpha_1 D \qquad (9)$$

where D is a dummy variable taking the value zero before the turning point, but unity afterwards. The introduction of D into eq.(9) obviously ensures a ratchet effect. The problem with (9) is that the implied time path for $\ln U_t^n$ is discontinuous. Replacing (9) by

$$\ln U_t^n = \alpha_0 + \beta_1 \ln U_t^w + \alpha_1 D + \beta_2 C \ln U_t^w \qquad (10)$$

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cov($lnU_t^{\,w}$, ε_t) is non-zero. This covariance is seen to be $p^* = \rho_{nw}\sigma_n\sigma_w - \beta\sigma_w^2$. If ρ_{nw} is zero, then the OLS estimator of β will be downward biased; if hypercyclicality is found using OLS it is likely to be an underestimate if anything. However, shocks are likely to affect both white and non-white unemployment in the same way, hence ρ_{nw} is likely to be positive (and near to unity) in such a simultaneous process. Even so, we would still expect p^* to be negative, because σ_n and σ_w are likely to be of the same order of magnitude. So $\rho_{nw} > 0$ would tend to make any potential OLS bias smaller, but we would still expect to underestimate the hypercyclical parameter β.

where $C\ln U_t^n$ is an interactive dummy with C= 1 only when unemployment is rising, serving two purposes. Firstly, the introduction of β_2 in eq.(10) permits the hypercyclicality parameter β in eq.(8) to vary over the cycle. Secondly, the time path of $\ln U_t^n$ will be continuous provided the appropriate restriction is imposed at a spline point coinciding with the $\ln U^w$ turning point. For example, if the spline point $\ln \tilde{U}^w$ is at a point of maximum $\tilde{U}^w$, then the restriction $\beta_2 \ln \tilde{U}^w = \alpha_1$ must be imposed to ensure continuity. On the other hand, if the spline point $\ln \tilde{U}^w$ is at a point of minimum $\tilde{U}^w$, the restriction $\beta_2 \ln \tilde{U}^w = -\alpha_1$ is necessary. Note that $\tilde{U}^w$ is a known value of U^w. It will later be shown how eq.(10) can cope with a sequence of turning points.

Eq. (10) captures a ratchet quite simply. What it comes down to is this. Suppose that we are at the bottom of a business cycle (where unemployment is high) and white unemployment moves in a downward direction. Let white unemployment fall by r log points until the top of the cycle is reached. Eq.(10) says that $\ln U_t^n$ will fall by $\beta_1 r$ log points. Now let the process reverse, with white unemployment rising by r log points. Non-white unemployment will rise by $(\beta_1 + \beta_2)r$ log points. Overall a gap of β_2 r log points will emerge between white and non-white rates over this hypothetical cycle. So $\beta_2 > 0$ indicates an upward ratchet effect.

IV Exploring the data

a Stationarity

It is evident from Figs 1 and 2 that there is a seasonal pattern in the data. However, it is also evident that the seasonal components are small relative to cyclical or trend components. In view of this, we considered it appropriate not to pay undue attention to the underlying seasonal process in the subsequent analysis . There is as yet, despite a large and developing literature, no clear consensus as to how the issue of seasonality in time series should be addressed.[8] We, therefore, applied a simple moving average filter = $(1+L+L^2+L^3)\ln U_t/4$ to all eight time series. If the seasonality is stochastic, this effectively filters out any seasonal variation in unemployment that may or may not be present. However, it does imply a specific seasonal process.[9]

We are aware that conducting econometric analyses on data that have been subject to any adjustments raises mis-specification issues, see Sims (1974), Wallis (1974, 1978) and more recently Ghysels and Perron (1993). However, all the results are repeated using the seasonally unadjusted data (but with deterministic seasonal dummies). The findings are virtually identical, emphasising the fact, because the seasonality is relatively mild, the raw and filtered data are not that much different.

Using this seasonally adjusted data, we then undertook tests for stochastic stationarity for all eight time series to establish for each unemployment series (logged values) whether it is trend stationary I(0)) or difference stationary I(1) .[10] In the case of the

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[8] See Harvey and Scott (1994).

[9] If the repeated stochastic seasonal pattern follows $(1 - L^4)\ln U_t = \varepsilon_t$, this can be written as $(1 - L)(1 + L)(1 - iL)(1 + iL)\ln U_t = \varepsilon_t$, where $i^2 = -1$. This implies four unit roots at $L = \pm 1$ and $L = \pm i$. This can be written as $(1 - L)F_{1t} = \varepsilon_t$, where $F_{1t} = (1 + L)(1 - iL)(1 + iL)\ln U_t$ is the moving average filter (multiplied by four).

[10] We test $(1-L)F_{it} = \alpha_1 F_{1t-1} + \varepsilon_t$ augmented by significant lagged values of the first difference of the regressand. The lag length is determined by maximising Akaike's information criterion (AIC). The critical values for the ADF test statistic are taken from Fuller (1976, P373). This is a special case of the (HEGY) equation developed by Hylleberg et al (1990),

$$(1-L^4)\ln U_t = \alpha_1(1+L)(1+L^2)\ln U_{t-1} - \alpha_2(1-L)(1+L^2)\ln U_{t-1} - \alpha_3(1-L)(1+L)\ln U_{t-2} - \alpha_4(1-L)(1+L)\ln U_{t-1} + \psi_1\Delta_4\ln U_{t-1} + ... + \psi_k\Delta_4\ln U_{t-k} + \gamma T + \delta_1 S_1 + \delta_2 S_2 + \delta_3 S_3 + \psi_0 + \varepsilon_t.$$

US data we took advantage of the fact that the time series commenced in 1954 in conducting these tests. The US data are all I(0) while the British data are I(1) at conventional significance levels.

So this simple unit root test comparison suggests a consistent difference between the British and US unemployment data. It is also not that surprising in that US unemployment appears to have a strong natural unemployment rate `attractor' with little or no tendency for upward drift. The British data over this period appears to have no equivalent natural rate attractor - the natural unemployment rate appears to have been subject to an upward drift over this period, reflecting the inflexibility of the labor market.[11] Nevertheless, given that unemployment rates, unlike most other time series data, have a lower and upper bound, one would have a reasonable prior that unemployment data should be I(0). With an I(1) process one would expect the upper or lower bound to be hit eventually even if it was a driftless random walk. So it is probably best to think of the British data exhibiting a pattern consistent with an I(1) process over this data period only.[12]

As the unit root tests suggest that unemployment rates for the British white and non-white data are I(1), we adopted the Engle and Granger (1987) two step procedure to test for cointegration for the British data only. The British non-white male

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when $\alpha_2 = \alpha_3 = \alpha_4 = 0$ implying seasonal unit roots at L = -1 and $L = \pm i$. (T is a deterministic trend and the S_i are seasonal dummies). Estimating a full HEGY equation, we found evidence of seasonal unit roots at a bi-annual and annual frequency, but there was no consistent pattern across all data series. Lindley (1999) explores the seasonality issue further.

[11] See Layard *et al.* (1991).

[12] Consider the root ϕ associated with time series $X_t = \phi X_{t-1} + u_t$. A useful interpretation is to think of this as not fixed but as a random coefficient. Thus over certain sub-periods this could be = 1, but < 1 at others. Taylor and Smith (1998) adopt this flexible approach. Also one needs to bear in

unemployment series cointegrate with white males as do non-white and white females.[13] This suggests that a unique underlying relationship exists between non-white and white British unemployment rates. Thus we feel justified in exploring underlying relationships such as eq.(8). With the US data the question of cointegration obviously does not arise, since the US series are all I(0).

b Hypercyclicality and ratchet effects

If unemployment is hypercyclical for non-whites, one would expect β to be greater than unity in eq.(8). OLS regression results are provided in Table 1, where a quadratic trend term has been included to capture non-cyclical effects. Thus we estimate

$$\ln U_t^n = \alpha + \beta \ln U_t^w + \gamma_1 t + \gamma_2 t^2 + e_t \tag{11}$$

where $\alpha = \ln \dfrac{U^{*n}}{(U^{*w})^{\beta}}$ from eq.(8). It is immediately apparent that there are strong differences between the US and British (both for males and females) results in two key respects.

First, non-white unemployment appears to be hypercyclical in Britain with a coefficient significantly greater than one attached to the white unemployment term

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mind the lack of power of all such unit root tests, that is there is a high probability of falsely accepting the null hypothesis. See Nelson and Plosser (1982) and Perron (1988).

13 For $\hat{e}_t = \ln U_t^n - \hat{\alpha} - \hat{\beta} \ln U_t^w - \hat{\gamma}_1 t - \hat{\gamma}_2 t^2$, we find ADF values of -4.69 and -2.95. The auxiliary regressions include significant lagged values of the first difference of the regressand. The lag length is determined by maximising AIC. The critical value of -2.89 is taken from MacKinnon (1991).

for both males and females, whereas in the US data there is no evidence of hypercyclicality - indeed the opposite appears to be the case. The US coefficient is well below unity in both cases. Secondly, the linear and quadratic trend terms (intended to capture non-cyclical movements upwards or downwards in relative unemployment) appear to work in opposite directions in the two countries. Differentiating eq.(11) with respect to time gives

$$d\ln U_t^n/dt \; - \; \beta \ln U_t^w/dt \; = \; \gamma_1 \; + \; 2\gamma_2 t \tag{12}$$

The term $\gamma_1 + 2\gamma_2 t$ measures the rate of convergency of non-white and white unemployment, abstracting from cyclical variations and allowing for any possible hypercyclical effects. Negative values, of course, indicate divergency. The rate of convergency/divergency becomes zero when $t = -\gamma_1/2\gamma_2$.

Since $\gamma_1 > 0$ and $\gamma_2 < 0$ in the US, eq.(11) implies that non-white and white unemployment diverge during the early part of the period but that this divergence reaches a maximum after roughly half the observation period, (52.7 quarters for males and 54.2 quarters for females). During the second half of the period, however, non-white and white unemployment rates converge again. The maximum long term divergency is 0.21 log points, implying that non-white unemployment was roughly 23% higher than white unemployment at that time. For females the maximum divergency was very similar, 0.22 log points. One explanation for convergence in the second half might be the `prison population' effect - the disproportionate increase in non-white prisoners starts to pull down rates. On the other hand a similar convergence is found in female rates, where there the prison population effect would be less strong. The non-white female prison population in 1998 was around 100 thousand compared with 890 thousand unemployed. Britain, incidentally, had only 791 non-white female prisoners in April 1998.

Britain had exactly the opposite experience with $\gamma_1 < 0$ and $\gamma_2 > 0$. Thus, during the early part of the observation period, non-white and white unemployment actually converge, but with convergency reaching a minimum very quickly for males after just 25 quarters. Afterwards non-white and white rates diverge markedly. By the end of the period, long-run divergency adds 0.23 log points to non-white unemployment, compared with a negligible effect for male non-whites in the US.

For females, divergency between non-white and white unemployment begins later - after 50 quarters. By the end of the period the divergency had added only 0.041 log points to non-white female unemployment. So this provides some evidence that the British labor market was somewhat more flexible for females. It is interesting to note that employment among non-white females was the fastest growing British group with the numbers economically active trebling from 214 thousand to 624 thousand over the period, despite apparent discrimination.

Overall, Table 1 confirms the previous more casual observations that things were better for British non-whites in the beginning but were a lot worse at the end and this deterioration affected British non-white males more quickly than British non-white females.

[INSERT TABLE 1 ABOUT HERE]

In Table 1, while diagnostic tests indicate that the residuals from eq.(11) are normally distributed, they are also autocorrelated and heteroscedastic. This is likely to be the result of dynamic mis-specification. To resolve these problems, we consider short-run variations in unemployment rates about an underlying relationship between non-white and white rates that is given by eq.(11).

By way of example, suppose the short-run relationship between U_t^n and U_t^w is

$$\ln U_t^n = b_0 + b_1 \ln U_t^w + b_2 \ln U_{t-1}^w + \mu \ln U_{t-1}^n + \varepsilon_t \tag{13}$$

Current values of U^n and U^w may not satisfy the underlying relationship (11) so we define a disequilibrium error

$$e_t = \ln U_t^n - \alpha - \beta \ln U_t^w - \gamma_1 t - \gamma_2 t^2 \tag{14}$$

We can now reparameterise eq.(13) as

$$\Delta \ln U_t^n = b_1 \Delta \ln U_t^w - \lambda e_{t-1} + \varepsilon_t \tag{15}$$

where $\beta = (b_1 + b_2)/\lambda$, $\alpha = b_0/\lambda$ and $\lambda = 1 - \mu$.

Equation (15) is a first order error correction model and its parameters having a clear interpretation. β is a the underlying elasticity of non-white unemployment with respect to white unemployment, whereas b_1 is the short-run elasticity. Both β and b_1 are hypercyclicality parameters. It is acceptable to save the fitted residuals from eq.(11) and estimate a higher order error correction model of the form

$$\Delta \ln U^n_t = b_1 \Delta \ln U^w_t + b_2 \Delta \ln U^w_{t-1} + b_3 \Delta \ln U^w_{t-2} + b_4 \Delta \ln U^w_{t-3} + b_5 \Delta \ln U^w_{t-4}$$

$$+ b_6 \Delta \ln U^w_{t-5} + c_1 \Delta \ln U^n_{t-1} + c_2 \Delta \ln U^n_{t-2} + c_3 \Delta \ln U^n_{t-3} + c_4 \Delta \ln U^n_{t-4}$$

$$+ c_5\Delta \ln U^w_{t-5} - \lambda \hat{e}_{t-1} + \varepsilon_t \qquad (16)$$

where $\hat{e}_{t-1} = \ln U^n_{t-1} - \hat{\alpha} - \hat{\beta}\ln U^w_{t-1} - \hat{\gamma}_1 t - \hat{\gamma}_2 t^2$. Since the two British difference stationary unemployment time series are cointegrated, the short-term disequilibrium relationship between them can be expressed in the above error correction forms. OLS estimation results for eq.(16) are provided in Table 2. Because the corresponding US series are stationary, strictly speaking the US short-run relationship does not need to be in error-correction form. However, for the sake of comparison an error-correction representation is presented in Table 2. The diagnostics now show that the residuals are normally distributed, non-autocorrelated and homoscedastic. As the disequilibrium term is taken from eq.(11), the estimates of the underlying parameters in eq.(16) are those provided in Table 1.

[INSERT TABLE 2 ABOUT HERE]

Table 2 also demonstrates strong differences between the US and British data. With the US data there is no evidence of hypercyclicality. However, with the British data, the size of the b_1 coefficients indicates that non-white unemployment displays hypercyclicality not only in the long-run but also in the *short-run* .

For the US, Table 2 shows that the short-run elasticities of non-white unemployment with respect to white unemployment are only 0.68 for males and 0.35 for females. Estimated adjustment parameters of 0.055 and 0.043 suggest that, for US males and females, only 5.5% and 4.3% of the disequilibrium in any period is eliminated in the following period

For the British data, the short-run elasticities of non-white unemployment with respect to white unemployment are estimated as 1.21 for males and 1.44 for females. These are significantly higher than unity and indicate short-run hypercyclicality. Estimated adjustment parameters suggest that, for British males and females, around 13% and 6.7% of any disequilibrium in one period will be eliminated in the following period. Thus adjustment towards the underlying relationship appears to be more rapid in Britain than in the US.

[TABLE 3 ABOUT HERE]

Now let us use the estimated coefficients of Table 1 to answer the following counterfactual question. Suppose Britain had enjoyed US white unemployment rates, what would its predicted non-white unemployment rate be? We can ask a corresponding question for the US, namely what is predicted non-white unemployment if it had the much worse British white rates of unemployment. Table 3 presents these results, where we have split these predictions into two equal periods (1971 (3) - 1985(1) and 1985(2) - 1998(2)) as well as presenting an overall figure. Comparing block (1) with (2) in Table 3, we see that US rates would have been considerably higher, but this effect is more marked in the second period. Similarly comparing (3) and (4) we see that British rates would have been considerably lower, and once again this effect is most marked in the second period.

But we can compare in a different way, with a slightly different interpretation. If we compare (3) with (2), then (2) now can be interpreted as telling us what British rates would have been if Britain had the same coefficients as the US. Once again the two period dichotomy is apparent (particularly for males). At first British rates would have been better, then they deteriorate in the second period. Similarly, we can compare (1) with (4), US rates would have been lower had they had the same coefficients as Britain.

c The spline function approach

An alternative way of capturing long run non-cyclical changes in relative unemployment rates is to model ratchet type effects by estimating spline functions similar to eq.(10). Generalising eq.(10) yields

$$\ln U_t^n = \alpha_0 + \beta_1 \ln U_t^w + \sum \alpha_i D_{it} + \beta_2 C_t \ln U_t^w \qquad (17)$$

where $C_t = 1$ when white unemployment is rising and $C_t = 0$ otherwise, and where

$$D_{it} = 1 \quad for \quad i = 1, 2, 3, \ldots\ldots\ldots, k\text{-}1 \qquad (18)$$

after the kth turning point in the U^w_t series but where all the D_{it} are all zero before. Thus the dummies switch on sequentially; that is, once the first turning point is reached, $D_{1t} = 1$ and stays on, once the second turning point is reached $D_{2t} = 1$ and stays on, etc.

The dummy variables in eq.(17) play roles similar to the dummy variable in eq.(10) and allow the possibility of ratchet effects. The C dummy implies that the hypercyclicality parameter β can vary over the cycle. In fact $\beta = \beta_1 + \beta_2$ when $C_t = 1$ and white unemployment is rising but $\beta = \beta_1$ when $C_t = 0$ and white unemployment is falling. If $\beta_2 > 0$ then β is greater in the recession than it is in the recovery and this is the ultimate factor behind any ratchet effects.

To ensure continuity in the $\ln U_t^n$ series, it is necessary to impose the restrictions $\beta_2 \ln \tilde{U}_i^w = -\alpha_i$ at the ith turning point of white unemployment (if it is a peak) and the restrictions $\beta_2 \ln \tilde{U}_i^w = \alpha_i$ at the ith turning point (if it is a trough). After imposing the restrictions eq.(17) becomes

$$\ln U_t^n = \alpha_0 + \beta_3 Z_{it} + \beta_1 \ln U_t^w + \beta_2 C_t \ln U_t^w \qquad (19)$$

where

$$Z_{it} = -\ln \tilde{U}_1^w D_{1t} + \ln \tilde{U}_2^w D_{2t} - \ln \tilde{U}_3^w D_{3t} + \ldots.. \qquad (20)$$

and the $\ln \tilde{U}_i^w$ are again known turning points (even numbers represent peaks and odd numbers represent troughs). To ensure complete continuity in the non-white unemployment series the restriction $\beta_2 = \beta_3$. must be imposed on eq.(19). Thus, the intercept dummies and any cyclical effects are combined into a single ratchet parameter β_2

OLS results for eq.(19) with $\beta_2 = \beta_3$ imposed are provided in Table 4. For US males the estimated value for β_2 is not significantly different from zero. This suggests that the $\beta = \beta_1 + \beta_2$ elasticity remains unchanged over the cycle. In addition, the β value appears to be less than unity. Therefore we find no evidence of full or partial hypercyclicality of male non-white unemployment with respect to white unemployment in the US. If fact, the estimated β suggests the reverse.

For female US data, we again find no evidence of hypercyclicality. However the estimated value β_2, is significantly different from zero, indicating that the β elasticity is greater in the recession than in the recovery. We, therefore, find evidence of a statistically significant ratchet effect for females. However the effect is somewhat mild. At most (by comparing high and low unemployment points) it added 0.048 log points to non-white unemployment.

For British data the β_2 estimates are in both cases significantly different from zero, implying that the elasticities of non-white unemployment with respect to white unemployment vary over the cycle. When white unemployment is rising, elasticities are estimated as $\beta = 1.21$ for males and $\beta = 0.88$ for females. However, when white employment is falling, the corresponding elasticities are 1.04 and 0.75. Clearly there is no evidence of hypercyclicality in the female data but the value of $\beta = 1.21$ in recessions, which is significantly different from unity, implies that non-white unemployment is partially hypercyclical with respect to white unemployment. The estimated value $\beta = 1.04$, obtained when white unemployment is falling, is not significantly different from unity.

For Britain strong ratchet effects are found for both male and female non-white unemployment. The ratchet coefficients, β_2 are highly significantly different from zero. For the male data the ratchet effects adds roughly 0.246 log points to non-white unemployment over the full observation period. This is a very powerful ratchet effect. For females the ratchet adds roughly 0.166 log points. But notice that this alternative way of viewing the data no longer supports hypercyclicality among females. So the spline approach seems to give more emphasis to the ratchet and less to hypercyclicality whereas the time trend approach gave more emphasis to hypercyclicality.

In Table 4 the restriction $\beta_2 = \beta_3$ has been imposed on eq.(19). F-tests indicate that the restriction is data-acceptable for both US equations. For Britain the restriction is data-acceptable for female data and very close to being data-acceptable for male data. The restricted and unrestricted versions of eq.(19) in fact yield very similar qualitative findings.

[INSERT TABLE 4]

As in Table 1, the diagnostic statistics in Table 4 indicate autocorrelated, heteroscedastic but normally distributed residuals. Again this is likely to be the result of dynamic mis-specification and we therefore estimate short run relationships in error correction form.[14]

The ECM estimates are presented in Table 5 and it can be seen that the diagnostic statistics are no longer a problem. Estimates of long-run parameters for the ratchet models were given in Table 4. As in Table 2, it can be seen that there are also large differences between the US and British data as far as estimated short-run coefficients are concerned. Again, there is no evidence of short-run hypercyclicality for the US. The short-run elasticities of non-white unemployment with respect to white unemployment are estimated as 0.69 and 0.40 for males and females respectively. For British data, however, we again find short-run hypercyclicality. The estimated values of b_1 in Table 5 are 1.15 and 1.36, respectively. As in Table 2, the estimated λ coefficients suggest a more rapid adjustment to underlying relationships for the British data than for the US data.

[INSERT TABLE 5 ABOUT HERE]

V Concluding comments

[14] As before, both British series are found to cointegrate with ADF values of -4.64 for males and -3.42 for females.

This investigation of the unemployment records of British and US non-whites has demonstrated that the two labor markets accommodated these disadvantaged groups in very different ways. It provides confirmation of the view that the US labor market is more flexible than that of Britain, as evidenced by the fact that non-white unemployment is partially hypercyclical with respect to white unemployment in Britain, whereas no such pattern is found in the US data. In recessions, British non-whites are forced into unemployment to an extent that is not found in the US. There is fairly clear cut evidence that, in Britain, a ratchet effect operates that results, cycle by cycle, in a worsening gap between non-white and white unemployment in Britain. This ratchet appears to be stronger for male data than for female. In contrast there appears no evidence of long run divergence between male non-white and white unemployment in the US. In fact a declining gap between non-white and white males is apparent from the mid-eighties onward. The picture for US females is less clear, but again there appears to have been some improvement from the mid-eighties onward.

However, the flexible labor market brings disadvantages as well as advantages and it would be misleading to paint too pessimistic picture of British non-whites. Another major source of ethnic disadvantage is lower earnings for ostensibly similar qualifications. It is clear that greater flexibility in the US labor market has been accompanied by a widening of earnings differentials and real earnings for those in the lowest decile of the earnings distribution have fallen in real terms. Although earnings inequality has widened in the British labor market, this has not been as severe as the US. Real earnings have shown some growth at the lower tails of the distribution.[15] There are signs that earnings differentials for British non-whites, particularly those born in Britain who constitute about 30% of all non-whites, have been eroded to the extent that it is not possible to detect any white differential once characteristics (experience, qualifications and so on) are controlled for. So it seems to be a choice of `low pay and employment' - the US option or `better pay and unemployment' - the British option.

[15] See Goodman and Webb (1994) and Blackaby *et al.* (1997).

Table 1. UK OLS estimation results for eq.(11).

$$\ln U_t^n = \alpha + \beta \ln U_t^w + \gamma_1 t + \gamma_2 t^2 + \varepsilon_t$$

	US Males	US Females	British Males	British Females
Est of α	0.993 (0.000)	1.08 (0.000)	0.0345 (0.005)	0.577 (0.000)
Est of β	0.771 (0.000)	0.694 (0.000)	1.230 (0.000)	1.208 (0.000)
Est of γ_1 (*100)	0.797 (0.000)	0.817 (0.000)	-0.190 (0.007)	-1.721 (0.000)
Est of γ_2 (*10000)	-0.755 (0.000)	-0.753 (0.002)	0.374 (0.000)	1.629 (0.000)
R Squared	0.93	0.94	0.98	0.87
Serial Correlation[a]	93.11 (0.000)	93.20 (0.00)	88.97 (0.000)	96.19 (0.000)
Normality[b]	2.67 (0.262)	1.53 (0.464)	0.42 (0.807)	2.39 (0.302)
Heteroscedasticity[c]	0.130 (0.718)	1.31 (0.251)	12.20 (0.000)	3.44 (0.064)

Notes; *Using the filtered data, from 1971 (3) to 1998 (2).*
P values in parentheses.
a. Lagrange Multiplier test.
b. Test for Skewness and Kurtosis of the residuals.
c. Regression of squared residuals on squared fitted values.

Table 2. An error-correction approach to modelling non-white/white unemployment.

$$\Delta \ln U_t^n = b_1\Delta\ln U_t^w + b_2\Delta\ln U^w_{t-1} + b_3\Delta\ln U^w_{t-2} + b_4\Delta\ln U^w_{t-3} + b_5\Delta\ln U^w_{t-4} + b_6\Delta\ln U^w_{t-5} + c_1\Delta\ln U^n_{t-1} + c_2\Delta\ln U^n_{t-2} + c_3\Delta\ln U^n_{t-3} + c_4\Delta\ln U^n_{t-4} + c_5\Delta\ln U^n_{t-5} - \lambda\hat{e}_{t-1} + \varepsilon_t$$

	US Males	US Females	British Males	British Females
Est of b_1	0.681 (0.000)	0.348 (0.000)	1.21 (0.000)	1.44 (0.000)
Est of b_2	-0.240 (0.029)	*	-0.83 (0.000)	-1.04 (0.000)
Est of b_3	*	*	*	*
Est of b_4	-0.316 (0.003)	*	*	-0.03 (0.035)
Est of b_5	0.538 (0.000)	*	*	0.56 (0.000)
Est of b_6	-0.256 (0.011)	*	*	-0.421 (0.001)
Est of c_1	0.547 (0.000)	0.536 (0.000)	0.702 (0.000)	0.737 (0.000)
Est of c_2	-0.205 (0.029)	*	*	*
Est of c_3	0.328 (0.001)	*	*	*
Est of c_4	-0.427 (0.00)	-0.318 (0.000)	*	-0.409 (0.000)
Est of c_5	0.188 (0.040)	0.283 (0.000)	*	0.335 (0.000)
Est of λ	0.056 (0.013)	0.043 (0.029)	0.130 (0.000)	0.067 (0.000)
R Squared	0.89	0.79	0.94	0.93
Serial Correlation[a]	0.627 (0.428)	0.445 (0.504)	3.80 (0.051)	2.32 (0.127)
Normality[b]	0.447 (0.799)	0.268 (0.874)	5.18 (0.075)	1.53 (0.464)
Heteroscedasticity[c]	0.079 (0.778)	1.95 (0.162)	0.551 (0.458)	12.21 (0.000)

Notes; *Using the filtered data, from 1971 (3) to 1998 (2)*
P values in parentheses.
** denotes statistically insignificant and is therefore dropped..*
a. Lagrange Multiplier test.
b. Test for Skewness and Kurtosis of the residuals.
c. Regression of squared residuals on squared fitted values.

Table 3. Predicting values for the level of non-white unemployment.

	Average for Males	Average for Females
(1)		
Predicted Level of US Non-White Unemployment.	11.5 overall 11.8 period 1 11.1 period 2	11.8 overall 12.9 period 1 10.7 period 2
(2)		
Predicted Level of US Non-White Unemployment using British Unemployment	14.6 overall 12.4 period 1 17.3 period 2	13.3 overall 12.8 period 1 14.2 period 2
(3)		
Predicted Level of British Non-White Unemployment	13.1 overall 9.4 period 1 18.5 period 2	14.3 overall 12.7 period 1 16.1 period 2
(4)		
Predicted Level of British Non-White Unemployment Using US Unemployment	9.2 overall 8.7 period 1 9.1 period 2	11.4 overall 13.1 period 1 9.8 period 2

Note Period 1 = 1971(3) - 1985(1)
Period 2 = 1985(2) - 1998(2)

Table 4. OLS regression results for the restricted eq.(19) .

$$\ln U_t^n = \alpha_0 + \beta_1 \ln U_t^w + \beta_2(Ct \ln U_t^w + Z_{it}) + \varepsilon_t$$

	US Males	US Females	British Males	British Females
Est of α_0	0.886 (0.000)	0.889 (0.000)	0.259 (0.436)	1.011 (0.000)
Est of β_1	0.937 (0.000)	0.862 (0.000)	1.037 (0.000)	0.751 (0.000)
Est of β_2	-0.004 (0.717)	0.063 (0.002)	0.175 (0.000)	0.126 (0.001)
R Squared	0.89	0.83	0.98	0.82
Serial Correlation[a]	95.17 (0.00)	98.97 (0.000)	92.60 (0.00)	100.22 (0.000)
Normality[b]	2.169 (0.338)	7.31 (0.026)	2.42 (0.298)	2.59 (0.240)
Heteroscedasticity[c]	0.014 (0.907)	8.37 (0.002)	13.36 (0.000)	4.55 (0.024)

Notes; ***Using the filtered data, from 1971 (3) to 1998 (2).***
P Values in parentheses.
a. Lagrange Multiplier test.
b. Test for Skewness and Kurtosis of the residuals.
c. Regression of squared residuals on squared fitted values.

Table 5. An error-correction approach to modelling the ratchet effect in unemployment.

$$\Delta \ln U_t^n = b_1\Delta\ln U^w_t + b_2\Delta\ln U^w_{t-1} + b_3\Delta\ln U^w_{t-2} + b_4\Delta\ln U^w_{t-3} + b_5\Delta\ln U^w_{t-4} + b_6\Delta\ln U^w_{t-5}$$
$$+ c_1\Delta\ln U^n_{t-1} + c_2\Delta\ln U^n_{t-2} + c_3\Delta\ln U^n_{t-3} + c_4\Delta\ln U^n_{t-4} + c_5\Delta\ln U^n_{t-5} - \lambda\hat{\nu}_{t-1} + \varepsilon_t$$

	US Males	US Females	British Males	British Females
Est of b_1	0.669 (0.000)	0.395 (0.000)	1.148 (0.000)	1.364 (0.000)
Est of b_2	-0.231(0.039)	*	-0.840 (0.000)	-1.034 (0.000)
Est of b_3	*	*	-0.221 (0.050)	*
Est of b_4	-0.316 (0.003)	*	*	*
Est of b_5	0.535 (0.000)	0.283 (0.000)	0.535 (0.000)	0.570 (0.000)
Est of b_6	-0.249 (0.014)	*	-0.354 (0.003)	-0.418 (0.001)
Est of c_1	0.556 (0.000)	0.336 (0.000)	0.721 (0.000)	0.745 (0.000)
Est of c_2	-0.207 (0.029)	*	0.207 (0.032)	*
Est of c_3	0.333 (0.001)	*	*	*
Est of c_4	-0.432 (0.000)	-0.286 (0.000)	-0.457 (0.000)	-0.410 (0.000)
Est of c_5	0.200 (0.030)	*	0.310 (0.002)	0.336 (0.000)
Est of λ	0.044(0.044)	0.019 (0.309)	0.105 (0.003)	0.050 (0.001)
R Squared	0.90	0.79	0.96	0.93
Serial Correlation[a]	0.502 (0.479)	1.917(0.166)	0.261 (0.609)	2.103 (0.147)
Normality[b]	0.833 (0.659)	0.931 (0.628)	6.001(0.050)	1.233 (0.540)
Heteroscedasticity[c]	0.078 (0.780)	1.206 (0.272)	0.369 (0.543)	12.939 (0.000)

Notes; *Using the filtered data, from 1971 (3) to 1998 (2)*

a. Lagrange Multiplier test.
b. Test for Skewness and Kurtosis of the residuals.
c. Regression of squared residuals on squared fitted values.

P values in parentheses.
** denotes statistically insignificant and is therefore. dropped.*

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