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Abstract

This paper analyses the stochastic properties of UK nominal and real wages over the period 1750-2015 using fractional integration techniques. Both the original series and logged ones are analysed. The results generally suggest that nominal wages exhibit a higher degree of persistence, which reflects relatively long lags between inflation and wage adjustments. Endogenous break tests are also carried out and various structural breaks are identified in both series. On the whole the corresponding subsample estimates imply an increase over time in the degree of persistence of both series.

JEL-Codes: C120, C130, C220, J300.

Keywords: nominal and real wages, mean reversion, persistence, fractional integration, structural breaks.

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1. Introduction

This paper provides new evidence on the stochastic properties (more specifically, mean reversion and persistence) of both nominal and real wages in the UK over a long time span, i.e. from 1715 to 2015, using fractional integration techniques. In a previous study based on a much longer sample (more precisely, from 1260 to 1994) Gil-Alana (2005) had focused exclusively on the long run and only in the case of real wages and found evidence of a unit root; Caporale and Gil-Alana (2006) had instead analysed the same data applying a procedure due to Robinson (1994) which tests for the presence of unit (and fractional) roots at both the zero and the cyclical frequencies. They concluded that the former play a more significant role, even when allowing for a break in 1875, which coincides with the beginning of the Second Industrial Revolution and the move to "New" unionism representing workers in a wider set of industries and resulting in a mass labour movement.

The present study makes a twofold contribution compared to the two mentioned above. Specifically, it uses a shorter sample starting a few decades before the period over which the Industrial Revolution (i.e. the transition to new manufacturing processes) took place (approximately from 1760 to around 1820 or 1840) to examine the behaviour of both nominal and real wages over a relatively more homogeneous period, and it allows for multiple breaks rather than a single one.[1] The first extension is essential because the degree of nominal and real wage inertia or persistence can be very different: in the former case it depends on the extent to which changes in current prices or inflation feed through

[1] Another interesting issue is how wage persistence can be explained by permanent worker, employer, and match heterogeneity. Carneiro et al. (2022) analyse it in the case of Portugal in the context of a dynamic panel model by applying a bias correction to deal with the incidental parameter problem. They report that the uncorrected estimates understate wage persistence and overstate the importance of permanent unobserved heterogeneity in driving it. The present study focuses instead on a comparison between persistence in real and nominal wages respectively using long runs of data to examine long-memory properties in the case of the UK.

to nominal wages in the same period (with a stronger response normally occurring when prices increase rather than decrease), whilst in the latter case it is determined by possible labour market frictions, a slow and/or only partial adjustment taking place when real wages or the mark-up of prices over marginal costs are not very responsive to demand pressures. In other words, the speed at which nominal or real wages adjust in response to exogenous shocks is mainly caused by labour market rigidities rather than the nature of the shocks themselves.

Concerning the second extension of the analysis, i.e. the tests for multiple endogenous breaks, this type of investigation is very important because a number of different wage determination regimes (e.g., before and after the creation of trade unions, or different periods in the unionisation movement) have been in place during the period examined, which should be taken into account empirically. Their existence also implies that different theoretical models might be appropriate for explaining wage behaviour in different sub-samples; for instance, in the more recent decades 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, has been found to describe well the UK experience (see Layard et al., 1991, 1994). Similarly, different policies might be required in different time periods, the general consensus being that supply side policies, such as wage bargaining reforms, are most effective in reducing unemployment (see Layard et al., 1994), whilst demand management policies do not have permanent effects (see Barrell et al., 1994).

The layout of the paper is as follows. Section 2 outlines the econometric methodology. Section 3 describes the data and presents the empirical results. Section 4 offers some concluding remarks.

2. Methodology

For our purposes we use fractional integration methods that have the advantage of being more general and flexible than standard ones based on the unit root versus stationarity dichotomy that only allows for integer degrees of differentiation.

The chosen specification, which also includes deterministic terms (namely a constant and a linear time trend), is the following:

$$y(t) \;=\; \alpha \;+\; \beta\, t \;+\; x(t); \qquad (1 \;-\; L)^{d}\, x(t) \;=\; u(t). \tag{1}$$

where y(t) is the observed time series; α an β are unknown coefficients on the constant and the linear time trend respectively; L is the lag operator, i.e. $L^k x(t) = x(t-k)$, and x(t) is assumed to be integrated of order d, or I(d), where d is another parameter to be estimated from the data. Finally, u(t) is a I(0) or short-memory process which is assumed in turn to be a white noise and to exhibit (weak) autocorrelation.

The estimation is carried out using a Whittle function in the frequency domain as in the fractional integration tests of Robinson (1994), which are widely applied in the empirical literature. In addition, the Bai and Perron's (2003) tests re used to detect any possible structural breaks.

3. Data Description and Empirical Results

The data examined are nominal and real wages in the UK at an annual frequency over the period from 1715 to 2015. They have been constructed by the Bank of England and are available from https://ourworldindata.org/grapher/nominal-wages-consumer-prices-and-real-wages-in-the-uk-since-1750. Both series, whether in their raw or logged form, exhibit an increasing trend (see Figure 1 and 2).

Figures 1 and 2 about here

Table 1 reports the estimates obtained when using the original data. It can be seen that in the case of nominal wages neither the intercept nor the time trend are statistically significant regardless of the assumption made about the disturbances; by contrast, the intercept is significant in the case of real wages. The estimates of d are higher than 1 in all cases (namely for both nominal and real wages and with both white noise and autocorrelated residuals), the I(1) hypothesis always being rejected in favour of $d > 1$.

Tables 1 and 2 about here

Table 2 reports the results for the log-transformed data. The estimates of d are now smaller compared to those based on the original data. More precisely, in the case of white noise disturbances (panel i) they are equal to 1.50 for nominal wages and 0.99 for real wages, and the I(1) hypothesis is rejected in favour of higher values of d for nominal wages but not for real ones. In addition, a significant positive time trend coefficient is found for the logged real wages. When allowing for autocorrelation (panel ii), the estimates of d are smaller (1.34 for nominal wages and 0.80 for real wages) and, as in the white noise case, the I(1) hypothesis is rejected for nominal wages but not for real ones. The time trend coefficient is now significant for both series and nominal wages have a higher slope coefficient.

On the whole our findings suggest that nominal wages exhibit a higher degree of persistence than real ones, i.e. they are characterised by higher (lower) rigidity (flexibility) compared to the latter; this is a similar result to what is normally found for the US, where real wage flexibility and nominal wage rigidity are thought to reflect relatively long lags between inflation and wage adjustments (see, e.g., Branson and Rotenberg, 1979, and Sachs, 1980); by contrast, in the case of most other European countries the opposite normally holds in the presence of inflationary shocks, which is a consequence of a relatively high degree of indexation of wages to prices (and a relatively low degree of inertia in the determination of nominal wages – see, e.g., Coe, 1985 and Arpaia and Pickelmann, 2007).

Next we focus on the log-transformed data and also test for structural breaks in this case; [2] in particular, using the Bai and Perron (2003) approach we detect five breaks in the case of nominal wages (in 1792, 1835, 1879, 1924 and 1969), and four in that of real ones (in 1858, 1888, 1935 and 1975 – see Table 3). These broadly correspond to well known historical events or policy measure, more specifically: (i) in the case of nominal wages approximately to the start of the war between Britain and revolutionary France, the Poor Law amendment that tightened relief, the start of the Anglo-Zulu war, the first Labour government (headed by Ramsay MacDonald), and the White Paper "In Place of Strife" issued by the Labour Goverment to reform the Trades Union movement; (ii) in the case of real wages approximately to the Indian Mutiny, the Convention of Constantinople guaranteeeing free maritime passage through the Suez canal in war and peace, the first "two-day weekend" giving workers extra time off instead of making redundancies, and the coming into force of the Equal Pay Act and Sex Discrimination Act.

Table 4 reports the estimates of d for each of the corresponding subsamples, six for nominal wages and five for real ones. In the former case, the time trend coefficient is found to be positive and significant in the first subsample (1750 – 1792) as well as the last two, namely 1925 – 1969 and 1970 – 2015, in the latter its value being much higher. The estimates of d show a mononotic increase over the first four subsamples (until 1924), decrease slightly during the fifth one, and increase again during the last one. It is noteworthy that all of them are significantly above 1, which implies a rejection of the I(1)

[2] The test results were essentially the same when using the original series; they are not reported for brevity's sake but are available from the authors upon request.

hypothesis, the only exception being the first subsample for which the estimate of d is below 1 and thus mean reversion is found.

Tables 3 and 4 about here

As for logged real wages, the trend is now found to be positive for the third (1889 – 1935) and the fourth (1936 – 1975) subsamples; in addition, the I(1) hypothesis cannot be rejected for the first four subsamples whilst it is rejected in favour of I(d, d > 1) for the last one.

However, these results should be taken with caution given the relatively small number of observations for each subsample the very wide confidence bands for the values of d. Thus, in what follows, we restrict the number of breaks in the series to two, specifically in 1835 and 1924 for logged nominal wages, and in 1738 and 1967 for logged real ones (Table 5).

Tables 5 and 6 about here

Table 6 displays the corresponding subsample estimates for d. It can be seen that in the case of logged nominal wages the time trend is no longer significant in any subsample, whilst the estimates of d are 1.07 (for 1750 – 1835), 1.79 (for 1836 – 1924) and 1.54 (for 1925 – 2015), and the unit root null hypothesis cannot be rejected for the first subsample, whilst it is in favour of d > 1 for the remaining two. As for logged real wages, the time trend is significant in the last two subsamples (1874 – 1967) and (1968 – 2015), its coefficient being much higher in the latter. The estimates of d increase monotonically, from 0.84 in the first subsample to 0.95 in the second one and 1.36 in the last one, the null hypothesis of I(1) not being rejected in the first two cases.

4. Conclusions

This paper has analysed the stochastic properties of UK nominal and real wages over the period 1750-2015 using fractional integration techniques that are more general than standard approaches restricting the differencing parameter to be an integer. Endogenous break tests have also been carried out since different wage determination regimes have been in place over the time period considered. Nominal and real wage developments matter because they have implications for price stability and competitiveness at country level. The results generally suggest that nominal wages exhibit a higher degree of persistence, and thus adjust with relatively long lags to inflation shocks. Also, on the whole the subsample estimates imply an increase over time in the degree of persistence of both series.

The fact that nominal wages exhibit a higher degree of persistence than real ones indicates that the UK labour market is more similar to the US one than those of the other European countries. In the latter set of economies real wage rigidity versus nominal wage flexibility is frequently found as a result of a relatively high degree of indexation of wages to prices. Persistence in real wages affects international competitiveness negatively since it implies that labour (and thus production) costs do not adjust quickly in response to shocks; in such cases appropriate labour market policies should be designed to increase flexibility and restore competitiveness. It would appear that the UK has generally had a competitive advantage given the higher degree of flexibility of its labour market.

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Table 1: Estimated coefficients using the original data

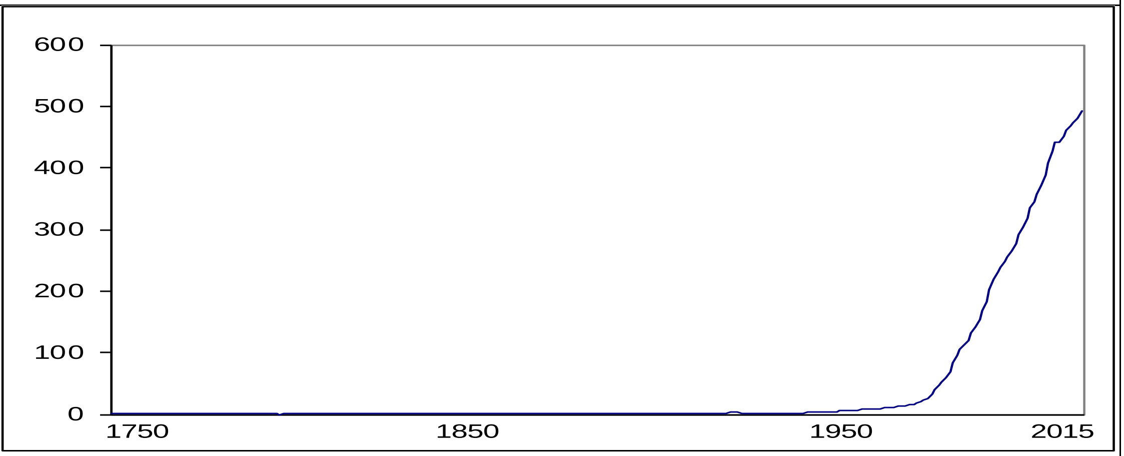
i) No autocorrelation (white noise) errors			
	d (95% band)	Intercept (t-value)	Time trend (t-value)
Nominal wages	1.70 (1.63, 1.80)	---	---
Real wages	1.41 (1.32, 1.53)	43.1323 (12.86)	---
ii) With autocorrelation (Bloomfield) errors			
	d (95% band)	Intercept (t-value)	Time trend (t-value)
Nominal wages	1.60 (1.53, 1.70)	---	---
Real wages	1.30 (1.21, 1.43)	43.1889 (12.59)	---

Table 2: Estimated coefficients using the logged transformed data

i) No autocorrelation (white noise) errors			
	d (95% band)	Intercept (t-value)	Time trend (t-value)
Nominal wages	1.50 (1.39, 1.63)	-1.2514 (-35.43)	---
Real wages	0.99 (0.91, 1.13)	3.7588 (67.15)	0.0092 (2.81)
ii) With autocorrelation (Bloomfield) errors			
	d (95% band)	Intercept (t-value)	Time trend (t-value)
Nominal wages	1.34 (1.22, 1.49)	-1.2518 (-33.89)	0.0251 (3.67)
Real wages	0.80 (0.73, 0.88)	3.7257 (70.57)	0.0092 (7.59)

Figure 1: Time series plots (original data)

Nominal wages



Real wages

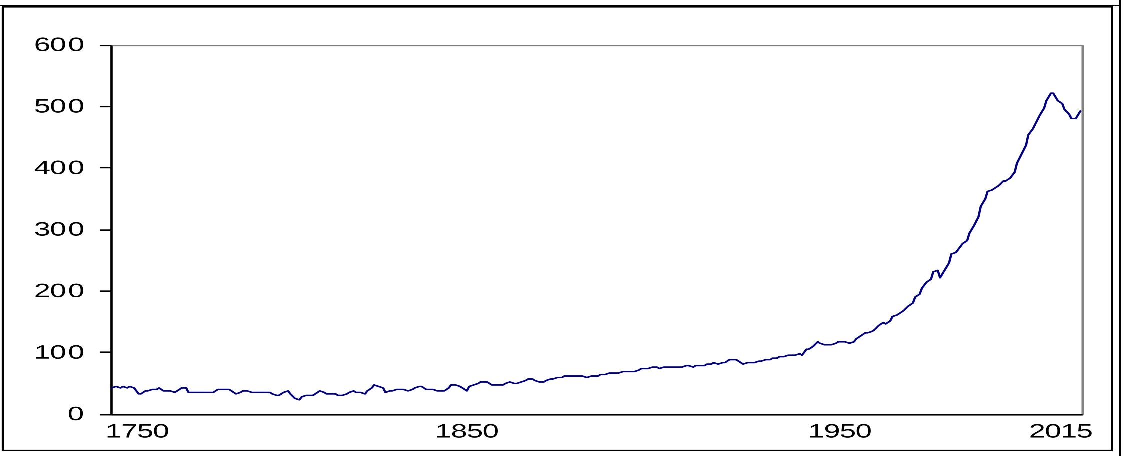
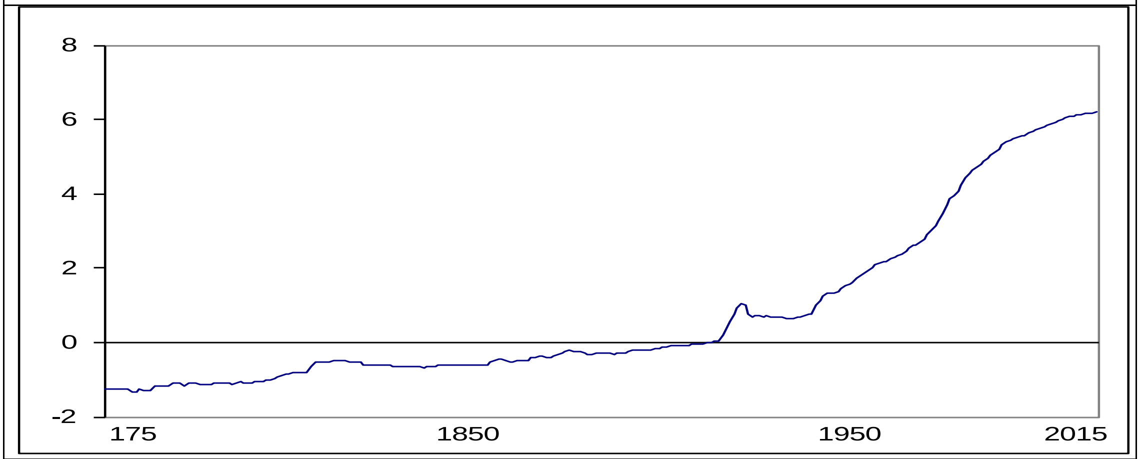


Figure 2: Time series plots (logged data)

Nominal wages



Real wages

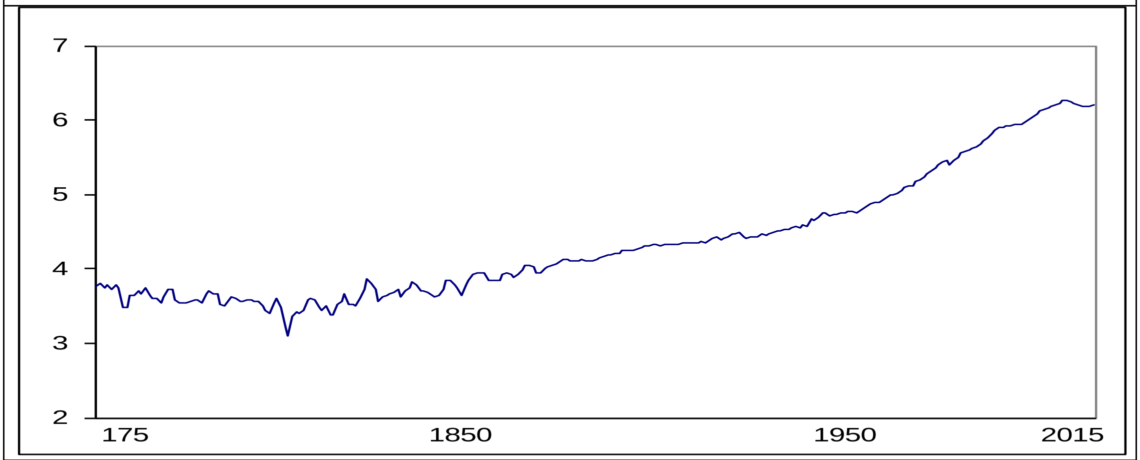


Table 3: Bai and Perron's (2003) results with multiple breaks

Series	N. of breaks	Break dates
Log of nominal wages	5	1792, 1835, 1879, 1924, 1969
Log of real wages	4	1858, 1888, 1935, 1975

Table 4: Estimated coefficients using a model with structural breaks

i) Logged nominal wages			
	d (95% band)	Intercept (t-value)	Time trend (t-value)
1750 – 1792	0.52 (0.30, 0.83)	-1.2713 (-52.49)	0.00523 (4.83)
1793 - 1835	1.26 (1.06, 1.57)	-1.0400 (-30.91)	---
1836 - 1879	1.50 (1.08, 2.00)	-0.6629 (-33.76)	---
1880 - 1924	1.82 (1.47, 2.35)	-0.3124 (-7.34)	---
1925 – 1969	1.30 (1.14, 1.54)	0.6982 (16.77)	0.03865 (2.22)
1970 – 2015	1.67 (1.51, 2.13)	2.7967 (102.64)	0.11105 (3.88)
i) Logged real wages			
1750 – 1858	0.82 (0.64, 1.16)	3.7612 (48.16)	---
1859 – 1888	1.07 (0.59, 2.09)	3.9415 (145.61)	---
1889 – 1935	0.68 (0.38, 1.02)	4.2098 (248.99)	0.00692 (7.36)
1936 – 1975	0.96 (0.81, 1.18)	4.5442 (176.07)	0.02228 (6.21)
1976 – 2015	1.52 (1.29, 1.90)	5.4567 (318.07)	---

Table 5: Bai and Perron's (2003) results with the number of breaks equal to 2

Series	N. of breaks	Break dates
Log of nominal wages	2	1835, 1924
Log of real wages	2	1738, 1967

Table 6: Estimated coefficients using a model with two structural breaks

i) Logged nominal wages			
	d (95% band)	Intercept (t-value)	Time trend (t-value)
1750 – 1835	1.07 (0.95, 1.24)	-1.2519 (-36.22)	---
1836 – 1924	1.79 (1.50, 2.16)	-0.6400 (-19.20)	---
1925 - 2015	1.54 (1.43, 1.70)	0.7238 (22.19)	---
i) Logged real wages			
1750 – 1873	0.84 (0.69, 1.15)	3.7636 (50.20)	---
1874 – 1967	0.95 (0.85, 1.09)	4.1136 (189.71)	0.01070 (5.85)
1968 – 2015	1.36 (1.18, 1.64)	5.1386 (269.88)	0.02373 (2.49)