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The Long-Run Determinants of UK Wages, 1860–2004

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Abstract

As it is almost 50 years since Phillips (1958), we analyze an historical series on UK wages and their determinants. Huge changes have occurred over this long run, so congruence is hard to establish: real wages have risen more than 6 fold, and nominal 500 times; laws, technology, wealth distribution, and social structure are unrecognizably different from 1860. We investigate: wage rates and weekly earnings; real versus nominal wages; breaks over 1860–2004; non-linearities, including Phillips' non-linear response to unemployment; 'trade union power' and unemployment benefits; and measures of excess demand, where workers react more to inflation when it rises.

1 Introduction

Following the introduction by Phillips (1958, 1962) of the famous 'Phillips curve' relation between wage inflation and unemployment, many studies have investigated wage inflation in the UK: salient contributions include Dicks-Mireaux and Dow (1959), Lipsey (1960), Sargan (1964, 1980), Godley and Nordhaus (1972), Nickell (1990), and Layard, Nickell and Jackman (1991). Most of these approaches have treated the labor market as the source of excess demand for labor, which in turn influenced wage inflation, although other models postulated competition between employers and employees over the profit share. Initially, the specification was in terms of nominal wages, followed by models with real-wage equilibria, and finally inflation expectations were accorded a key role, becoming dominant in the 'new-Keynesian' approach to price inflation. Most UK quarterly, post-war models have considered price and wage inflation jointly with unemployment, where their interactions induced a 'wage-price spiral', still of concern to policy makers–see http://www.bankofengland.co.uk/publications/news/2007/003.htm. Once a major focus, that aspect has not been at the centre of recent attention, so we will reconsider it.

The model of price inflation developed in Hendry (2001) considered a plethora of economic theories for its determining variables: no theory was individually sufficient, and almost all helped. Historical contingencies, primarily major wars and oil crises, also played a major role in addition to the economic influences, 'modeled' by individual-year indicators (one-off, or impulse, dummies). The same eclectic approach will be adopted for wages here, allowing for a range of possible influences.

The structure of the paper is as follows. Section 2 discusses the data series to be analyzed, considers some important measurement issues, describes the treatment of turbulent events, and compares two measures of nominal wages, namely hourly wage rates and average weekly earnings. Next, sections 3

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and 4 discuss the economic and econometric theories underlying our study, and section 5 replicates two major earlier empirical studies, by Phillips (1958) and Sargan (1964). Section 6 presents the model of UK wage inflation over this long run of data, 1860–2004, and investigates its properties. Finally, section 7 concludes. All the empirical analyses and graphics use *PcGive*, *Autometrics* (see Doornik, 2006, and Hendry and Doornik, 2006) and *PcGets* (see Hendry and Krolzig, 2001).

2 Measurement issues

2.1 Data series

The data set is annual for the UK over 1860–2004, and was derived from a number of sources, detailed in the appendix. The main series are constant-price GDP (denoted Y); prices (P, the implicit deflator of Y); nominal broad money (M); interest rates (Treasury-bill rate R_s, bond rate R_l: see Ericsson, Hendry and Prestwich, 1998); employment (L), unemployment (U) and working population (*Wpop*); nominal average weekly wage earnings (W: see Feinstein, 1990, building on Bowley, 1937, also see Crafts and Mills, 1994); nominal hourly wage rates (W_r: see Shadman-Mehta, 2000); normal hours (H); world prices (P_e); a trade union membership measure (*TU*); the 'replacement ratio' from unemployment benefits (B); and the nominal effective exchange rate (E). Other variables are constructed from this base set. Capital letters denote the original variables, and lower-case letters the corresponding logarithms (so $y = \log Y$). The estimation sample period is usually 1863–2004 after creating lags and allowing for missing data.

2.2 Measurement errors

An extended discussion of data accuracy is provided in Hendry (2001), who argues that the variables of interest are measured with sufficient accuracy to merit modeling despite the long historical period–which must entail substantial errors of measurement, both conceptual and numerical (many of the relevant components were not recorded at the time). Our working assumptions remain that: (a) the levels data are integrated once ($\mathsf{I}(1)$) with superimposed major breaks, so 'look' like $\mathsf{I}(2)$ series; and (b) measurements thereof have up to $\mathsf{I}(1)$ deviations from the desired theoretical counterparts. Consequently, wage inflation is treated as $\mathsf{I}(0)$ with breaks, but measured with an $\mathsf{I}(0)$ error. Mis-measurement in such processes therefore induces non-constancy in empirical models, but is consistent with the effects of revisions to post-war quarterly inflation time series (see Hendry, 1995, ch.14). Even in long non-stationary historical time series, despite including war years, constant-parameter equations are feasible (in the sense of Hendry, 1996: previous analyses have successfully modeled money demand–see, e.g., Ericsson *et al.*, 1998, and Escribano, 2004–and price inflation–Hendry, 2001), although measurement caveats must be borne in mind when interpreting any reported models.

2.3 Descriptive statistics

Figure 1, panel a, records nominal weekly wage earnings and price levels (in logs), with real wages and output per person employed in panel b, nominal changes in panel c, and real wage growth in panel d.[1] To establish the scale, note that the level of nominal average weekly earnings has increased *four hundred*

[1]Graph panels are lettered notionally as a, b, c, ..., row by row.

and fifty fold over the sample, and nominal wages per hour even more (*five hundred and seventy fold*). While the patterns of w and p are similar, the former has increased much faster, leading to real wage increases of between $5\frac{1}{2}$–7 fold. Wage inflation has varied from around $+25\%$ to worse than -20%, and again is similar to price inflation (panel c), with a few marked departures.[2] Such variations would swamp most I(0) measurement-error effects.

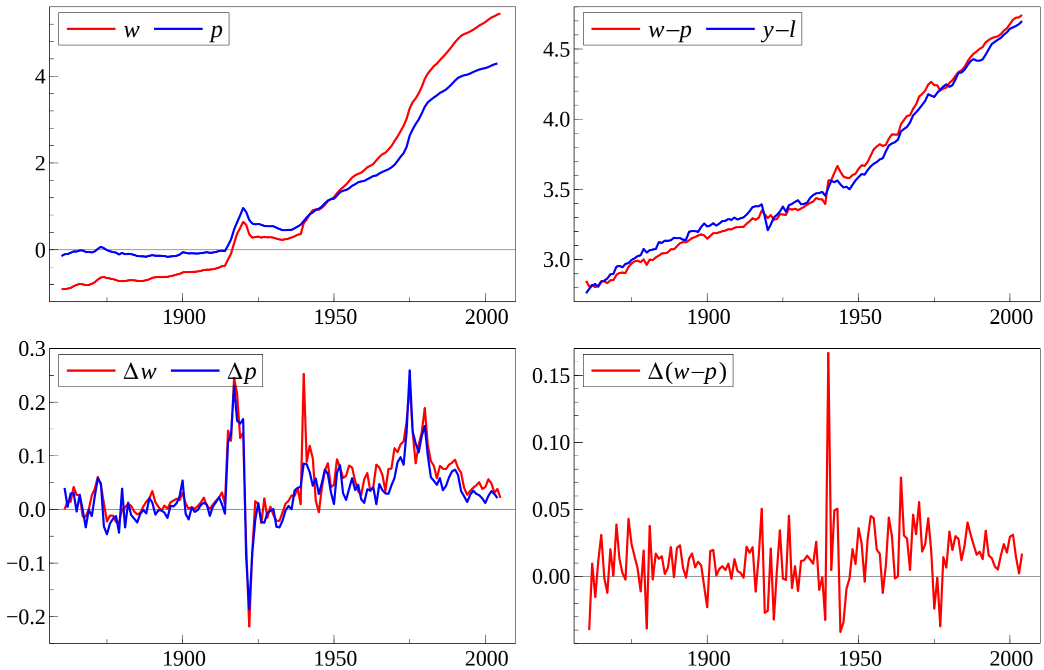

Figure 1 Nominal and real UK wages, prices, output per capita, and inflation rates.

The following descriptive statistics summarize means and standard deviations (SDs) of annual changes in log average weekly wages and wage rates respectively:

mean (SD)	1863–2004	pre-1914	1914–1945	pre-1945	post-1945
Δw	0.044 (0.061)	0.010 (0.018)	0.041 (0.097)	0.022 (0.063)	0.075 (0.042)
Δw_r	0.043 (0.067)	0.010 (0.025)	0.036 (0.107)	0.020 (0.070)	0.074 (0.046)
$\Delta(w-p)$	0.014 (0.024)	0.009 (0.015)	0.011 (0.038)	0.010 (0.026)	0.019 (0.020)
$\Delta(w_r-p)$	0.012 (0.029)	0.009 (0.020)	0.006 (0.050)	0.007 (0.035)	0.018 (0.018)

Table 1 Means and SDs of annual changes in log average weekly wages and wage rates.

Thus, mean annual nominal UK wage inflation was about 2.2% pre- and 7.5% post-1945 respectively, as the last two columns show, although the post-1945 SD is lower at 4.2%. Mean annual real-wage inflation was both much lower at 1.4% pa over the whole period, and more stable, with an SD of 2.4%. Despite such an apparently slow growth rate, the inexorable laws of compounding show that over the 140+ years, average real wages rose roughly 6 fold, more than half of that since 1945 (i.e., 3.2-fold), so individual prosperity in the UK is recent. This is consistent with the results on the rate of growth in real wages of 1.2% per year for 1813–1913 in Crafts and Mills (1994). A standard ‘popular’ perception is that the second half of the nineteenth century was a ‘Golden Age’ in the UK, which was the workshop

[2]Despite having lived through the large rises and falls during 1916–22, Keynes (1936) still argued wages were inflexible.

of the world where 'pax Britannica' ruled etc. In fact, the UK was at war in almost every year of that century; and GDP per head and its growth were both low by recent standards, albeit not by comparison with previous centuries. We conjecture that recorded perceptions of standards of living were formulated by the then rich (whose status was later eroded), who were the only historians of the time given the restricted availability of education in England till after WWII. However, as Crafts (1988) makes clear, detailed UK growth pre-1945 does not sustain many simple generalizations.

Conversely, there was a major social structural break in 1946, with the nationalization of many basic industries, introduction of a national health service, and improvements in state provision of education, unemployment insurance and pensions. Moreover, the Beveridge (1942, 1945) Reports essentially mandated UK Governments to keep a low level of unemployment using Keynesian policies, and the outturn of 1.5% on average over 1946–66 was historically unprecedented. Figure 2 reports the unemployment rate (a), its change (b), growth in output per capita as a measure of demand fluctuations (c), and excess demand for labor, U^d from Hendry (2001), in panel d.[3] There are four distinct epochs of unemployment, marked as shown in panel a, commencing with a distinct pattern of business cycle fluctuations, a jump in the mean and variance of U_r, the era of low mean and variance through post-war reconstruction, ending with a further jump in mean and variance from the late 1970s. If U_r is to be part of the explanation of Δw, then Δw would need to manifest such patterns conditional on other explanatory variables, either in the model of U_r or that of Δw.

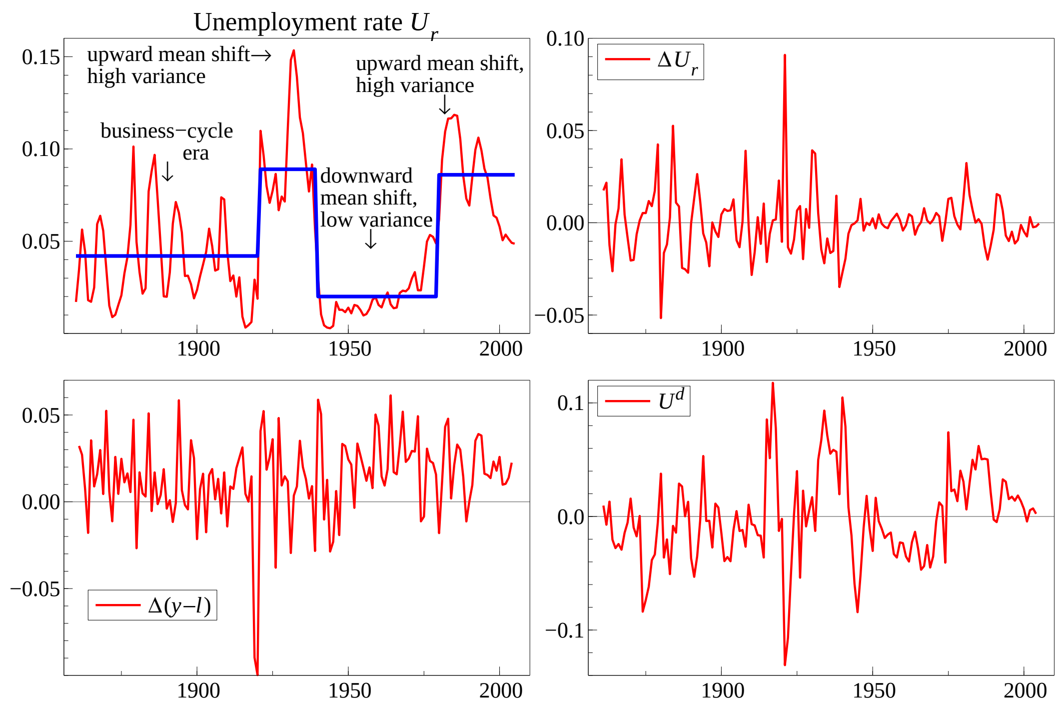

Figure 2 Unemployment rate, its change, growth in output per capita, and excess demand for labour.

As a baseline for judging later whole-sample fit, regressing Δw_t on two past values of itself and of past inflation yields $\widehat{\sigma} = 4.1\%$: thus, contemporaneous price inflation is a fundamental determinant of nominal wages since the SD of $\Delta(w-p)_t$ is just 2.4%. The standard neo-classical theory of wage determination is that the real marginal cost of labor (usually taken as the real wage) equilibrates to the real

[3] U^d captures disequilibrium unemployment based on steady-state growth, whereby the unemployment rate rises when the real interest rate exceeds the real growth rate ($R_l - \Delta p > \Delta y$) and vice versa.

marginal revenue product of labor. Given the evidence for a constant-returns technology at the aggregate level in Hendry (2001), then the ratio of real wages to average product should vary around a constant level. The similar trends in real wages and output per person employed in fig. 1b are consistent with that. The large increase in labor costs (particularly national insurance) should ensure $(w-p)$ grew more slowly than $(y-l)$, and their average growth rates were indeed 1.35% and 1.42% respectively.

Other than a spike in 1940, real wage growth shows little evidence of the turbulence present in its two nominal components (fig. 1d). This confirms the importance of conditioning on price inflation, noting that $\Delta(w-p)_t$ entails doing so with a coefficient of unity, and seems an excellent example of a co-breaking relationship (see Hendry and Massmann, 2007), an issue to which we now turn. We note the difficulty for an 'annual contract' model of wage bargaining, as workers could not possibly have correctly anticipated the major jumps and falls in inflation entailed by wars, and no econometric model seems able to forecast them. A simpler hypothesis is that of within-year adjustments at times of high inflation, and although some economists would interpret that as simultaneity, a simple interpretation is just as a 1–1 transformation of the joint density of w_t and p_t into the density of w_t and $(w-p)_t$. This is different from a conditional/marginal factorization, in that the latter may still depend on Δp_t, so the issue of valid conditioning remains. To check this hypothesis, we undertook impulse saturation on the conditional distributions, as explained in section 4.2.

2.4 Turbulent periods

Research on price inflation for this data set revealed an abundance of 'outliers' (22 residuals larger than $3\widehat{\sigma}$ in absolute value), matching the many economic policy and exchange-rate regime shifts, major wars, crises, and substantial legislative and technological changes. Such shifts are evident in both of the nominal series in fig. 1c, and have been recorded in related time series by numerous other authors (see e.g. Darne and Diebolt, 2004).

Indicators, or impulse dummies, correct for both innovation and additive 'outliers' in the regressand, and also adjust for any aberrations in regressors, so summarize deviations unaccounted for by the economic variables (see Salkever, 1976). In dynamic models, indicators have more complicated effects (see Doornik, Hendry and Nielsen, 1998), and even more complicated in cointegrated processes (Nielsen, 2004), but can be handled appropriately. If economic variables really matter in a model, then their effects should not be removed by also including indicators. Conversely, indicators can reduce correlations induced by chance matching of shifts. Finally, since zero residuals for observations removed by indicators can distort test statistics, indicators in an initial unrestricted model can be combined into an overall index to minimize such effects: see Hendry and Santos (2005). Section 4.2 discusses the new technique of indicator saturation developed by Hendry, Johansen and Santos (2004).

2.5 Measuring UK wage inflation

There are two main historical time series of wage-related indices, denoted w and w_r. The first is analyzed here, the second by Shadman-Mehta (1995), but we briefly compare these. The measures differ somewhat in their concept (earnings versus rates), and hence the impact of changes in hours and perhaps labor taxes thereon, but if these two series do not cointegrate with each other, they cannot cointegrate with the same determinants. Many theories of 'wage inflation' are imprecise about the precise measure, but empirically this matters greatly.

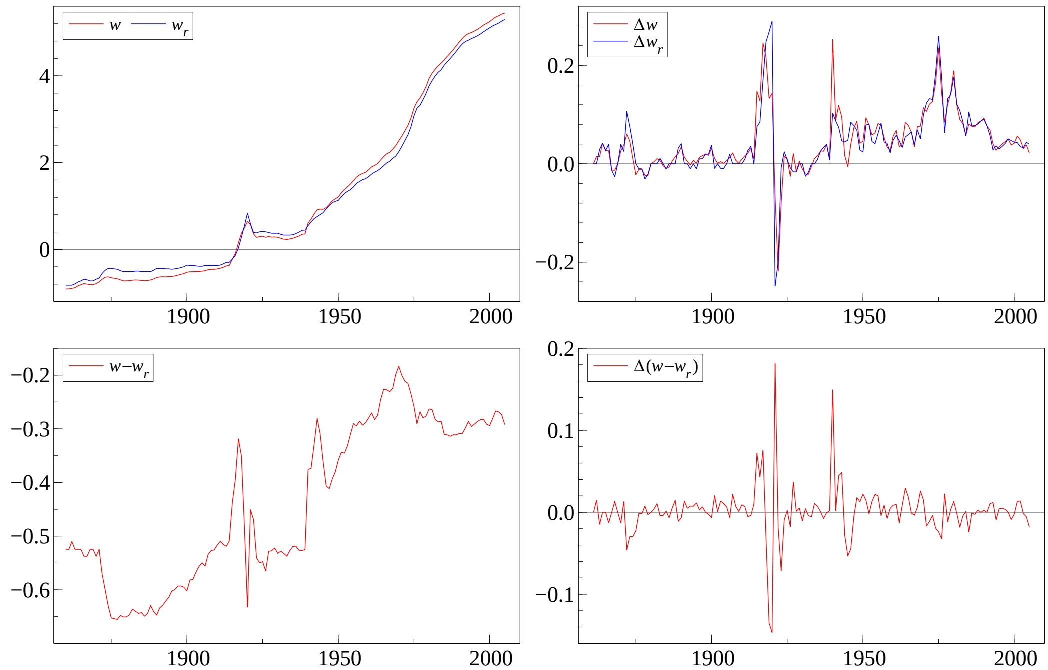

Figure 3 Two measures of UK wages, their inflation rates, and differentials.

First, figure 3, panel a, records the log-wage levels (adjusted to be zero in 1985), with their rates of change in panel b, the differential $(w - w_r)$ in panel c, and its growth in panel d. Although the levels and annual inflation rates look similar, the differential has a trend (roughly in line with the sparse data on falls in annual hours worked), and $\Delta\,(w - w_r)$ has a mean of 0.2% pa with a standard deviation of 3.2%, much larger than for $\Delta\,(w - p)$, but is dominated by the volatile periods 1914–1922 and 1939–1945, which look from panel b to be driven by a slightly different timing of peaks and troughs.

Cointegration was studied in a 2-variable VAR with 2 lags, an unrestricted constant, restricted trend, and unrestricted indicators for each year 1914–1922 inclusive and 1940 (to remove the main outliers) (see Johansen, 1995). The trace test (Tr(r) for the hypothesis $\mathsf{H}(r)$ of r cointegrating relations apparently rejects with the value 55.2^{**} and a cointegrating vector close to $(1, -1)$. However, both $\widehat{\alpha}_i > 0$, warning that there is no convergence. Thus, empirical models of one of these series need not be appropriate for the other. Henceforth we model w.

The residual standard errors from this VAR also provide an order-of-magnitude estimate of how well econometric equations might fit, namely, around $\widehat{\sigma} = 2.0\%$. This is similar to the value in section 2.4, and, given such evidence on data inaccuracy, suggests the final model has a respectable fit.

3 Economic theory relationships

There is a vast literature on the theory of wage determination, from the subsistence theory in Ricardo (1971, ch.5), through the classical Walrasian view of clearing labor markets with no 'involuntary' unemployment, the Keynesian focus on nominal rigidities, to the micro-founded models in the literatures on search and matching (see Pissarides, 2000), imperfect information, adverse selection and efficiency wages (see *inter alia* Weiss, 1990, and Shapiro and Stiglitz, 1984; Akerlof and Yellen, 2001, provide a review of the last). We briefly note theories of the long-run determinants (sub-section 3.1) and dynamic adjustment (sub-section 3.2).

3.1 Long-run determinants

Profit-maximizing competitive firms should only hire workers up to the point at which their marginal revenue product (MP) equals the marginal cost (MC). Firms should obviously pay workers as little as possible to maximize profits, subject to efficiency wage arguments, but demand for labor between firms bids up wages so long as MC<MR. With a Cobb–Douglas production function (broadly upheld by the evidence), marginal revenue is proportional to nominal average product, so lacking data on hours, the measure in our data is PY/L. Marginal cost comprises wages (W) as well as labor taxes (τ, including fringe benefits such as pensions, health care), where data on the latter are not available for the whole sample (see Alesina and Perotti, 1997, for evidence on labor taxation). Price deflation is usually by a CPI for wages and a PPI for output, which adds a 'wedge' to some wage determination models; but that is not relevant here as we only have the GDP deflator.

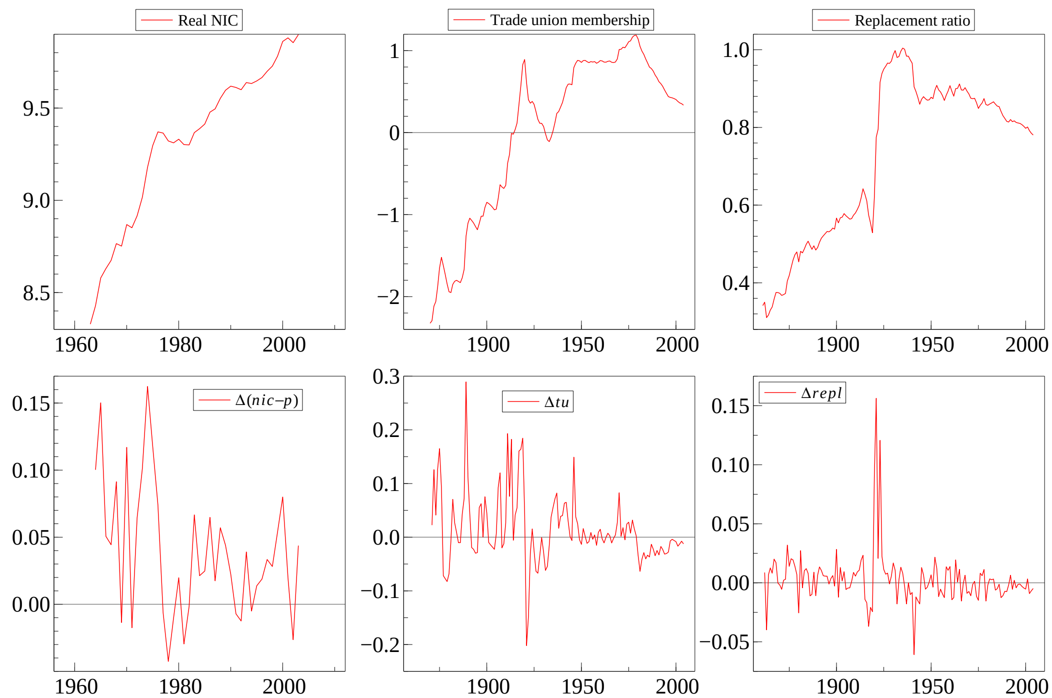

Figure 4 Real national insurance costs, trade union membership, replacement ratio, and their growth rates.

In general equilibrium, the effects of trade unions, unemployment benefits, employment legislation, and so on, mainly impact on unemployment rather than the real wage. For example, trade unions may raise wages at a given firm, but goods market competition then forces it to reduce its employment. Figure 4 records the available historical data in logs on real national insurance costs (from 1960: panel a), trade union membership (standardized: panel b), and real unemployment benefits (replacement ratio given by real benefits/lagged real wages: panel c), as well as their growth rates. All three series are highly non-stationary, and have increased greatly over the period whereas neither the level of unemployment nor the share of labor have trended (fig. 2: but see Manning, 1993). Thus, the long-run constant-price cointegration relationship in logs used here becomes simply:

$$w - p = y - l - \tau. \tag{1}$$

3.2 Dynamic adjustments

Dynamic adjustments have been explained by a range of factors, including staggered wage contracts, the role of trade unions (Oswald, 1985, provides a review), bargaining (see e.g., Lewis, 1986, MacDonald and Solow, 1981), the distinction between insiders and outsiders to the firm (see Oswald, 1993), institutional factors (see Calmfors and Driffill, 1988, and Soskice, 1990, for the UK), including labor laws (e.g., Nunziata, 2005) and expectations, as well as factors like indexation during the First World War. There is also a substantial literature on wage-price spirals, from early models such as Dicks-Mireaux and Dow (1959), where wage inflation was due to 'cost-push' in economies with price-setting firms through Phillips (1958), Sargan (1964), Blanchard (1987), Rowlatt (1988), and Surrey (1989) to Kolsrud and Nymoen (1998). Labor markets are certainly an important intermediary in the overall inflation process, since factor markets determine wages and the prices of capital goods. However, factor demands are derived from final demands, so the latter must be the direct determinants of price inflation. Unemployment (or that in relation to vacancies) is often the demand measure used, although capacity utilization and various measures of the output 'gap' also occur. The 'natural rate of unemployment' (see e.g., Friedman, 1977) or the related NAIRU, provide a widely-used framework. Nickell (1987) emphasizes the impact of the proportion of long-term unemployed on the Phillips curve. Macroeconomic factors such as the share of the exposed sectors of an open economy also could allow terms of trade shocks to have a significant effect: see Nymoen (1989). Analyses involving disequilibria from several sectors include Juselius (1992), and Metin (1995). Thus, we allow the data evidence to determine the adjustment process, including unemployment as a potential factor influencing the deviation in (1).

4 Econometric theory basis

Our approach is based on general-to-simple modeling. This, and the associated automatic modeling algorithms in PcGets and Autometrics, have been documented separately (see Hendry and Krolzig, 2005 and Doornik, 2007), and would take too much space to describe here. However, a few key features are noted in subsection 4.1, with three novel aspects, indicator saturation, collinearity and non-linearity, discussed in subsections 4.2–4.4 respectively.

4.1 General-to-simple modeling

General-to-simple modeling requires that all the candidate explanatory variables that might influence the variable being modelled are considered from the outset of an empirical analysis in a general unrestricted model (GUM). The costs of not doing so are either biased estimates from omitting relevant variables, or uncontrolled significance levels if variables are later added, or 'problems' corrected, in the light of adverse evidence. Key references establishing such difficulties include Anderson (1962, 1971), Mizon (1977), and Hendry (1979): Campos, Ericsson and Hendry (2005) summarize these developments. The converse costs are either large models with potentially diffuse estimates, or costs associated with model simplification to eliminate candidate regressors that happen not to matter. Over the last decade, there has been a revolution in the quality of algorithms for model selection, vastly improving over the negative findings in Lovell (1983) on 'data mining', where poor algorithms were at fault, through the multi-path search procedure in Hoover and Perez (1999), to the evidence of the excellent properties of PcGets in Hendry and Krolzig (2005) and Autometrics in Doornik (2007).

Two mistakes are always possible in any statistical inference or selection: false inclusion and false exclusion. We address these in turn.

Consider a candidate regressor set with $K = 100$ variables $\mathbf{x}_t$, all of which are in fact irrelevant, but all of which are included in a general model: we address the choice of K below, and how to handle $K > T$. At $\alpha = 1\%$ significance, corresponding to a $|\mathsf{t}|$ value of around 2.6 when the errors are approximately normal, $K\alpha = 1$ so 1 of the 100 would be retained by chance on average, with 99 out of the 100 correctly eliminated, leading to a vast increase in knowledge. Thus, it is relatively easy to eliminate irrelevant variables. Moreover, as almost no variables are retained, the estimated equation fit ($\widehat{\sigma}$) is not very different from that which would have been found had the correct state of nature been known at the outset, and no selection was required. Retaining one variable by chance sampling is a small cost for reducing $2^{100} \simeq 10^{30}$ possible models to one on average (which could be improved by joint testing). On 'conventional' calculations, however, the rejection frequency of the procedure (sometimes called its 'size') is $1 - (1 - 0.01)^{100} \simeq 0.63$, suggesting it performs badly.

Next, what happens to false exclusion? If k of the K candidates mattered, the 'power' to retain them would be close to that facing an investigator who used significance level α but knew precisely which k variables to enter, without any irrelevant candidates included. That implication follows analytically from the near unbiased estimate of the equation standard error σ and the elimination of almost all irrelevant variables, so essentially the same second-moment matrix of the $\mathbf{x}_t$ is used. Moreover, since the selection criterion is known ($|\mathsf{t}| > \mathsf{c}_\alpha$), any bias from only retaining coefficients when they are 'statistically significant' can be corrected, which also serves to attenuate the coefficient estimates on irrelevant variables that were retained by chance. Hendry and Krolzig (2005) document the theory and simulation evidence; and Hendry and Krolzig (2004) present an empirical application to growth regressions. Subsection 4.2 describes a procedure for removing outliers, such that normality is a reasonable approximation: and 4.3 discusses the impact of collinearity.

4.2 Indicator saturation

Both nominal and real wages fluctuated widely over the sample period, with 'extreme observations' frequently observed, often due to wars, technological and policy shifts. Such exogenous events induce structural breaks in the levels, which will manifest themselves as outliers in a differenced or cointegrated equation. A vast literature exists on identifying structural breaks at unknown dates, see *inter alia* Bai and Perron (1998) and Altissimo and Corradi (2003). Here, we implement indicator saturation as in Hendry *et al.* (2004).

Indicator saturation adds a complete set of impulse indicators $\{I_t,\ t = 1, \ldots, T\}$ to a model, but entered in feasible subsets to enable estimation. Hendry *et al.* (2004) establish the null distribution of the estimator of the mean in a location-scale IID distribution after adding T impulse indicators when the sample size is T. A two-step process is investigated: half the indicators are added, all significant indicators are recorded, then the other half is examined; finally the two significant subsets are combined. Indicators should be included as part of the GUM, alongside all potentially relevant variables to avoid spurious retention of either indicators or variables. This is especially important if non-linearities are present, since extreme observations due to non-linearities are observationally similar to outliers. As $K > T$ is certain to occur in this process, repeated combinatorial block entry will be needed, and α chosen to produce an acceptable value of $K\alpha$ (e.g., $\alpha = 0.1\%$ would yield $K\alpha = 1$ even for $K = 1000$): Castle and Hendry (2005) provide simulation evidence that the theory holds despite $K > T$ and

Autometrics (Doornik, 2007) provides automated software for $K >> T$.

We applied indicator saturation to the GUM outlined in section 6 using a 0.1% significance level. Various block sizes are automatically considered, but the choice of sample splits does not impact substantially on the results, matching the theoretical results in Hendry *et al.* (2004), who show the subset division should be irrelevant. Outliers were identified for the years 1918, 1940, 1942, 1943, 1975 and 1977. These dummies were included in the GUM in section 6 and selection was undertaken at a looser significance level (1%).[4] We also applied the indicator saturation to a GUM for nominal wages. Fewer indicators are retained for real wages compared to nominal (10 indicators were retained for Δw_t), suggesting that wages may co-break with prices (see Hendry and Massmann, 2007), which matches the data in fig. 1a, where the slope of nominal wages changes frequently over the sample period. The results suggest that there are some non-constancies, but the number of indicators retained is not large given the sample period.

4.3 Collinearity

Macroeconomic time series are generally both highly autocorrelated and intercorrelated. The former often reflects non-stationarity, which when it takes the form of unit-root, or integrated, processes, also induces the latter. Such a ‘problem’ is easily resolved by cointegration analysis and differencing. The presence of collinearity due to intercorrelation is pandemic in economics, and is often thought to favor parsimonious explanations. However, the bias-inducing impact of omitting relevant variables is largest when they are highly collinear with retained variables, whereas there is no impact of collinearity on rejection frequencies under the null. Worse still, omitting a relevant variable that experienced a shift in its data moments (as most have done) will induce non-constant parameters in the resulting model. Clements and Hendry (2005) show that one cannot avoid the consequences for forecasting by omitting a highly collinear relevant variable that experiences a location shift. Thus, the important issue is one of low power, due to the non-centrality of the distribution of a relevant variable’s coefficient being attenuated by collinearity. Over long sample periods with massive data variation and many structural breaks, such as that here, then collinearity does not seem to be the main difficulty confronting a cliometrician.

4.4 Non-linearity

Phillips identified the presence of non-linearities in the wage equation by modeling the change in the nominal wage rate as a function of the inverse of the unemployment rate. Our non-linear modeling framework comprises a test of functional form to first establish whether a reduction to linearity is viable, then non-linear modeling only if such a procedure is required. We discuss these two steps in turn.

4.4.1 Testing for non-linearity

Given the large number of potential explanatory variables and the collinearity between these, a useful test of non-linearity must handle high-dimensional parametrizations with highly collinear data. Castle and Hendry (2006) use a weighted function of all the regressors up to cubic, with weights given by the

[4] Some dummies were combined to form indexes (1975 and 1977) and a combined dummy for the Second World War augmented the two dummies found by indicator saturation (1942 and 1943) with two further dummies of opposite sign (1944 and 1945).

eigenvectors of the data variance-covariance matrix. Define the matrix of regressors as $\mathbf{X}' = (\mathbf{x}_1 \ldots \mathbf{x}_T)$, compute the eigenvectors $\mathbf{H}$ and eigenvalues $\mathbf{\Lambda}$ of $T^{-1}\mathbf{X}'\mathbf{X}$, and let:

$$\mathbf{z}_t = \mathbf{\Lambda}^{-1/2}\left[\left(\mathbf{H}'\mathbf{x}_t\right) - \overline{\left(\mathbf{H}'\mathbf{x}_t\right)}\right]. \tag{2}$$

We include both current and lagged variables, so $\mathbf{X}$ has dimension 20, but exclude any linear combination where $\lambda_i < 0.0001$ (contemporaneous inflation Δp_t is excluded from these regressors). Then (before including indicators) the GUM can be written as:

$$y_t = \beta_0 + \boldsymbol{\beta}_1'\mathbf{x}_t + \boldsymbol{\beta}_2'\mathbf{x}_{t-1} + \boldsymbol{\beta}_3'\mathbf{v}_t + \boldsymbol{\beta}_4'\mathbf{w}_t + u_t \tag{3}$$

where $\mathbf{v}_t = \left\{z_{i,t}^2\right\}$ and $\mathbf{w}_t = \left\{z_{i,t}^3\right\}$. The non-linearity test is of the significance of $\boldsymbol{\beta}_3 = \boldsymbol{\beta}_4 = \mathbf{0}$ in (3). For fixed regressors $\mathbf{x}_t$, the test is an exact F-test with $2n$ degrees of freedom for n elements in $\mathbf{z}_t$, so is correctly sized under the null, and has power against departures from linearity due to general quadratic or cubic effects (which can map to a smooth transition model with asymmetry and skewness: see Castle and Hendry, 2005). This test is a low-dimensional alternative to the test in White (1980), applicable for cubics and when $n(n+1)/2 > T$. This non-linear test is applied to the change in both nominal and real wages, where the linear regressors include y_{t-i}^d, U_{t-i}^d, $U_{r,t-i}$, Δpcl_{t-i}, $\Delta\left(y-l\right)_{t-i}$, $\Delta\left(ulc-p\right)_{t-i}$, $\Delta tu_{p,t-i}$, $\Delta\left(b-p\right)_{t-i}$, $\Delta repl_{t-i}$, for $i = 0, 1$ and Δp_{t-j} for $j = 1$. Two specifications are considered, namely inclusion of $\mathbf{v}_t$ and $\mathbf{w}_t$, and just inclusion of $\mathbf{w}_t$, as cubic functions preserve the signs of, but give more weight to, large deviations from equilibrium levels (see Hendry, 1984).

The results for the period 1872–2004 are reported in table 2.[5] Five linear combinations of the $\mathbf{X}$ matrix are excluded as $\lambda_i < 0.0001$. Hence, there are 15 degrees of freedom for the cubic test and 30 for the quadratic and cubic test. There is evidence of non-linearity in the real-wage equation, and that conclusion is robust to both quadratics and cubics, but there is no evidence of non-linearity in nominal wages. This suggests a functional form where non-linearity enters the wage equation through the coefficient of inflation. We anticipate an ogive form (e.g., LSTAR) with hysteresis: once the coefficient has increased, which can occur suddenly, it takes time to fall back, giving weight to the credibility arguments currently in vogue with the UK Monetary Policy Committee.

Variable	Test	F-test
$\Delta\left(w-p\right)$	$\mathbf{z}_t^2$,$\mathbf{z}_t^3$	$\mathsf{F}(30, 82) = 2.259^{**}$
$\Delta\left(w-p\right)$	$\mathbf{z}_t^3$	$\mathsf{F}(15, 97) = 2.269^{**}$
Δw	$\mathbf{z}_t^2$,$\mathbf{z}_t^3$	$\mathsf{F}(30, 82) = 1.494$
Δw	$\mathbf{z}_t^3$	$\mathsf{F}(15, 97) = 1.537$

Table 2 Non-linearity test results.

5 Replicating earlier studies

Two major historical UK studies are those by Phillips (1958) for the period prior to World War I, and by Sargan (1964) for the immediate post World War II era. We first seek to replicate these.

[5]Data on union power is only available from 1870 onwards.

5.1 Replicating the Phillips curve

A number of authors have replicated the original Phillips curve, both using his group-averaged data and the original annual observations: see Thomas (1984) and Desai (1975). Our own data set was updated to extend the sample back to 1860 to match Phillips (1958), using the hourly wage rates printed in Phelps-Brown and Hopkins (1950). We also used average weekly wage earnings, consistent with the analysis in section 6 and reported in fig. 5; both series yielded very similar replications of the Phillips curve.

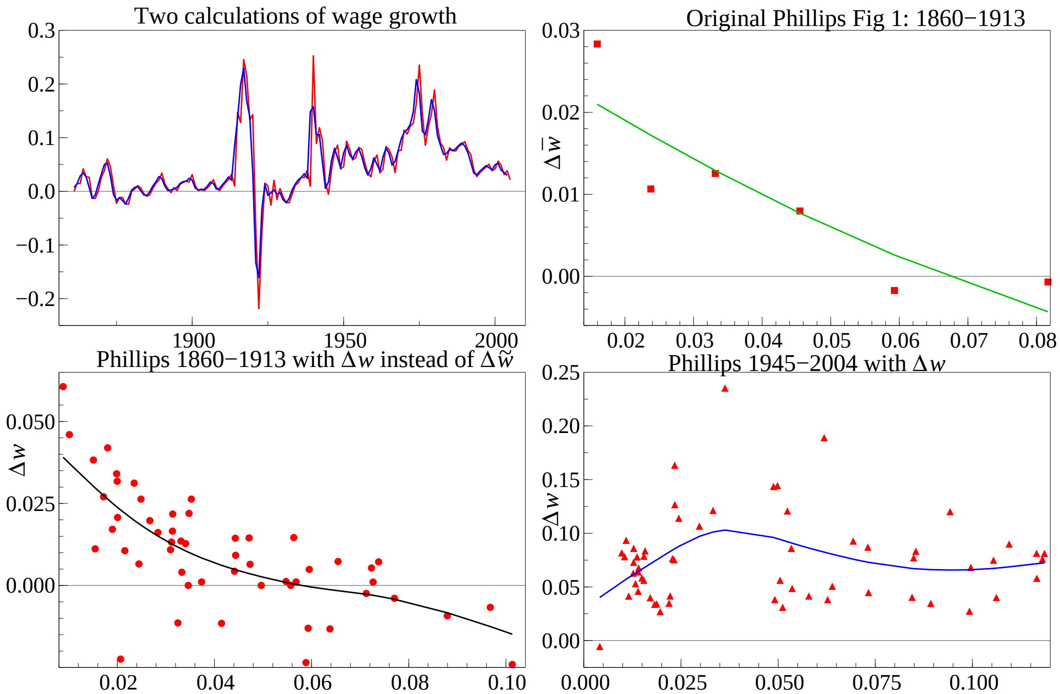

Figure 5 Growth rate of wages, Phillips curve replication, and Phillips curves using Δw.

There are two methods for calculating growth in the wage rate. The first is based on Phillips' definition, given by:

$$\Delta \widetilde{w}_t = \frac{0.5\,(W_{t+1} - W_{t-1})}{W_t} \tag{4}$$

and we compare this with the standard practice of taking the difference of the logs. Figure 5a records the two growth rates. There is a difference in timing at the peaks and troughs, but the correlation between the two series is 0.90.

	$\bar{U}_r$	$\Delta\bar{w}$
0-2	0.0160	0.0283
2-3	0.0238	0.0106
3-4	0.0333	0.0126
4-5	0.0455	0.0080
5-7	0.0589	-0.0017
7-11	0.0816	-0.0007

Table 3 Phillips curve replication.

We then replicate figure 1 in Phillips' original paper (panel b), where the fitted curve is based on the group averages, denoted $\Delta\bar{w}$ and $\bar{U}_r$, reported in table 3. Panel c shows the original data over the sample and a spline-based fit using Δw. Panel d shows the equivalent plot over 1945–2004, with the completely different picture of the well-known 'breakdown' of the Phillips curve.

Using modern methods and more extensive data, we improved a little over Phillips' pioneering results as follows ($\widehat{\mu}$ is the whole sample mean of $(w-p)-(y-l)$ of -1.85) for the sample $T = 1863$–1913:

$$
\begin{aligned}
\widehat{\Delta w_t} &= \underset{(.04)}{0.277}\ U_{r,t}^{-1} + \underset{(0.06)}{0.38}\ \Delta p_t - \underset{(0.033)}{0.055}\ \left((w-p)_{t-1} - (y-l)_t - \widehat{\mu}\right) \\
&\quad - \underset{(0.035)}{0.154}\ (R_l - \Delta p - \Delta y)_{t-1} \\
\mathsf{R}^2 &= 0.86^{\dagger}\ \ \widehat{\sigma} = 0.77\%\ \ \mathsf{SC} = -7.4\ \ \mathsf{V} = 0.18\ \ \mathsf{J} = 2.71^{**}\ \ \mathsf{F}_{\mathsf{reset}}(1,46) = 1.88 \\
\chi^2_{\mathsf{nd}}(2) &= 5.43\ \ \mathsf{F}_{\mathsf{ar}(1-2)}(2,45) = 2.51\ \ \mathsf{F}_{\mathsf{arch}(1-1)}(1,45) = 0.32\ \ \mathsf{F}_{\mathsf{het}}(8,38) = 0.23 \qquad (5)
\end{aligned}
$$

In (5), R^2 is the squared multiple correlation († denotes including a constant), $\widehat{\sigma}$ is the residual standard deviation, coefficient standard errors are shown in parentheses, *V* and *J* are the variance-change and joint parameter-constancy tests from Hansen (1992), and *SC* is the Schwarz criterion (see Schwarz, 1978). The diagnostic tests are of the form $\mathsf{F_j(k, T-l)}$ which denotes an approximate F-test against the alternative hypothesis j for: k^{th}-order serial correlation (F_{ar}: see Godfrey, 1978), k^{th}-order autoregressive conditional heteroskedasticity ($\mathsf{F}_{\mathsf{arch}}$: see Engle, 1982), heteroskedasticity ($\mathsf{F}_{\mathsf{het}}$: see White, 1980); the RESET test ($\mathsf{F}_{\mathsf{reset}}$: see Ramsey, 1969); parameter constancy ($\mathsf{F}_{\mathsf{Chow}}$: see Chow, 1960) over k^{th} periods; and a chi-square test for normality ($\chi^2_{\mathsf{nd}}(2)$: see Doornik and Hansen, 1994). Finally, * and ** denote significant at 5% and 1% respectively.

Thus, there is a strong inverse relationship to unemployment as Phillips found, small impacts from price inflation and the equilibrium correction of real wages to marginal product (denoted EqCM below), and a strong negative effect from the main determinant of U^d.

5.2 Replicating Sargan (1964)

We could replicate the quarterly wage equation in Sargan (1964) quite closely using annual data for Δw_t over the nearest sample equivalent of 1946–65 (see Hendry, 2003, for *average earnings*):[6]

$$
\begin{aligned}
\widehat{\Delta w_t} &= \underset{(0.13)}{0.64}\ \Delta p_t + \underset{(0.26)}{0.68} - \underset{(0.14)}{0.37}\ (w-p)_{t-1} + \underset{(0.003)}{0.009}\ t - \underset{(0.82)}{4.01}\ U_{r,t} \\
\mathsf{R}^2 &= 0.80\ \ \widehat{\sigma} = 1.01\%\ \ \mathsf{SC} = -8.7\ \ \mathsf{F}_{\mathsf{Chow(1964:1)}}(2,13) = 3.0 \\
\chi^2_{\mathsf{nd}}(2) &= 0.30\ \ \mathsf{F}_{\mathsf{ar}(1-4)}(4,11) = 0.36\ \ \mathsf{F}_{\mathsf{arch}(1-4)}(4,12) = 0.34\ \ \mathsf{T} = 1946\text{–}1965 \qquad (6)
\end{aligned}
$$

The major difference between using earnings rather than rates is that $U_{r,t}$ is significant. Sargan lacked good data on productivity, so used a linear deterministic trend as a proxy as in (6), but we have data on output per person employed, and using that instead yields:

$$
\widehat{\Delta w_t} = \underset{(0.12)}{0.52}\ \Delta p_t + \underset{(0.015)}{0.12} - \underset{(0.12)}{0.44}\ (w-p-y+l-\widehat{\mu})_{t-1} - \underset{(0.78)}{3.31}\ U_{r,t}
$$

[6] Recent revisions to the GDP deflator have resulted in a revised series over the entire post-WWII period due to methodological changes. Replicating the Sargan model with this data leads to the same specification with almost identical coefficients but two dummies are required to account for outliers.

$$\begin{aligned}
\mathsf{R}^2 &= 0.79 \;\; \widehat{\sigma} = 1.03\% \;\; \mathsf{SC} = -8.8 \;\; \mathsf{F}_{\mathsf{Chow(1956:1)}}(10,6) = 0.44 \\
\chi^2_{\mathsf{nd}}(2) &= 0.96 \;\; \mathsf{F}_{\mathsf{ar(1-4)}}(4,12) = 0.02 \;\; \mathsf{F}_{\mathsf{arch(1-4)}}(4,12) = 0.70 \;\; \chi^2_{\mathsf{het}}(6) = 2.8
\end{aligned} \tag{7}$$

where $\widehat{\mu}$ is again the whole sample mean (quite close to the subsample mean of -1.9). One also needs to take deviations from the mean of 5% for $U_{r,t}$ for the intercept to be interpretable, in which case it vanishes, consistent with the absence of any 'autonomous' wage inflation. There is a slight deterioration in fit, but letting $U_{rm,t} = U_{r,t} - 0.05$ delivers the simplified model:

$$\begin{aligned}
\widehat{\Delta w_t} &= \underset{(0.12)}{0.58} \; \Delta p_t - \underset{(0.12)}{0.48} \; (w - p - y + l - \widehat{\mu})_{t-1} - \underset{(0.24)}{1.95} \; U_{rm,t} \\
\mathsf{R}^2 &= 0.74 \;\; \widehat{\sigma} = 1.09\% \;\; \mathsf{SC} = -8.7 \;\; \mathsf{F}_{\mathsf{Chow(1956:1)}}(10,7) = 0.53 \\
\chi^2_{\mathsf{nd}}(2) &= 0.04 \;\; \mathsf{F}_{\mathsf{ar(1-4)}}(4,13) = 0.13 \;\; \mathsf{F}_{\mathsf{arch(1-4)}}(4,12) = 2.35 \;\; \mathsf{F}_{\mathsf{het}}(6,13) = 1.08
\end{aligned} \tag{8}$$

The long-run steady-state constant-inflation solution is that deviations of the wage share from its mean vary negatively with deviations of the unemployment rate from its mean of 5%. In the short run, real wages fall with inflation and unemployment, but are rapidly corrected by the wage share, which embodies the neo-classical proposition that the real wage equals the marginal product. However, instrumenting Δp_t by Δp_{t-1}, Δw_{t-1} and the constant yields a coefficient slightly larger than unity, so a real-wage reformulation is not precluded. Simultaneity would have induced an upward biased estimate for the coefficient of Δp_t in (8), so a 'measurement error' explanation is more plausible, possibly that Δp_t is a mis-measured proxy for expected inflation.

The findings from these equations are surprisingly consistent, though inconsistent with the direct replication of Phillips's work in fig. 5: although Phillips found no role for either price inflation or the EqCM from the wage share, we do here, making the relationship between the approaches that much closer. Indeed, simply enforcing the specification in (5) for the post-war period yields ($T =$1946–1965):

$$\begin{aligned}
\widehat{\Delta w_t} &= \underset{(0.13)}{0.70} \; \Delta p_t - \underset{(0.087)}{0.44} \; \left((w-p)_{t-1} - (y-l)_t - \widehat{\mu}\right) + \underset{(0.069)}{0.81} \; U^{-1}_{r,t} \\
\mathsf{R}^2 &= 0.81 \;\; \widehat{\sigma} = 0.95\% \;\; \mathsf{SC} = -9.0 \;\; \mathsf{V} = 0.09 \;\; \mathsf{J} = 0.72 \;\; \mathsf{F}_{\mathsf{reset}}(1,16) = 0.09 \\
\chi^2_{\mathsf{nd}}(2) &= 10.9^{**} \;\; \mathsf{F}_{\mathsf{ar(1-2)}}(2,15) = 0.01 \;\; \mathsf{F}_{\mathsf{arch(1-1)}}(1,15) = 0.22 \;\; \mathsf{F}_{\mathsf{het}}(6,10) = 0.26
\end{aligned} \tag{9}$$

The salient features of the specification in (8) reappear in (9), which also shares two significant effects with (5), albeit with different coefficients, the remaining two variables being insignificant. There is a larger coefficient on price inflation, which may reflect that past real-wage losses are also more important than in (5), but the stark difference from the model behind fig. 5 is muted. Interestingly, despite the much lower and less volatile unemployment rate in the post-war reconstruction era, the non-linear form $U^{-1}_{r,t}$ in (9) leads to a better fitting model than $U_{r,t}$ in (8).

6 Model of UK wage inflation 1860–2004

In keeping with the general-to-simple modeling philosophy outlined in section 4.1, the initial model of $\Delta(w-p)_t$ included the variables Δpcl_{t-i}, $\Delta repl_{t-i}$, and $(fna \times \Delta p)_{t-i}$ for $i = 0, 1$; $U_{r,t-j}$, $U^{-1}_{r,t-j}$ and $\Delta(y-l)_{t-j}$ for $j = 0, 1, 2$; as well as $(ulc - p - \widehat{\mu})_{t-j}$, U^d_{t-k} and $\Delta(w-p)_{t-k}$ for $k = 1, 2,$

where, as discussed following (11):[7]

$$fna = \frac{-1}{1 + 1000\Delta p_t^2}. \tag{10}$$

Four dummies were included based on the results of the impulse saturation applied in section 4.2, namely indicator variables for 1918, 1940, a combined dummy for 1975 and 1977 (first oil crisis and incomes policies), and a dummy for the shift during the Second World War of 3% up in 1942 and 43, followed by equal magnitude falls in 1944 and 45 (denoted *IWW2*). An intercept was also included. The unrestricted model yielded $\widehat{\sigma} = 1.20\%$ for 25 variables ($SC = -8.15$, and there were no significant mis-specification tests: without indicators $\widehat{\sigma} = 1.84\%$). The resulting specific model was selected using the PcGets conservative strategy, which delivers an overall significance level of approximately 1%. Including differences of the unemployment rate would result in perfect collinearity, but the final model suggested that $U_{r,t}$ and $U_{r,t-2}$ cancelled each other out: the restriction was accepted and imposed. Selection was then undertaken using the twice lagged difference.[8]

The resulting wage model is reported in equation (11), and is an updated version of that used in Hendry (2001), with a longer sample both before and after that research. Automatic model selection provided an important check on the previous model specification, and furthermore, the extended sample enabled *ex ante* forecasts to be computed given the same model specification. The $\mathsf{F}_{\mathsf{Chow}}(14, 105)$ test is for the 14 years from 1991 on, data not available in that earlier study.

$$\begin{aligned}
\Delta (w-p)_t = \; & \underset{(0.13)}{0.76}\; fna \times \Delta p_t + \underset{(0.002)}{0.010} + \underset{(0.05)}{0.39}\; \Delta (y-l)_t + \underset{(0.05)}{0.13}\; \Delta (y-l)_{t-2} \\
& - \underset{(0.01)}{0.08}\; (ulc - p - \widehat{\mu})_{t-2} - \underset{(0.04)}{0.14}\; \Delta_2 U_{r,t-1} + \underset{(0.013)}{0.027}\; I1918_t \\
& + \underset{(0.013)}{0.14}\; I1940_t - \underset{(0.009)}{0.05}\; I7577_t + \underset{(0.006)}{0.03}\; IWW2_t
\end{aligned} \tag{11}$$

$$\begin{aligned}
\mathsf{R}^2 &= 0.75 \;\; \widehat{\sigma} = 1.24\% \;\; SC = -8.51 \\
\chi^2_{\mathsf{nd}}(2) &= 3.0 \;\; \mathsf{F}_{\mathsf{ar}}(2, 130) = 0.18 \;\; \mathsf{F}_{\mathsf{arch}}(1, 130) = 0.40 \;\; \mathsf{F}_{\mathsf{reset}}(1, 104) = 0.01 \\
\mathsf{F}_{\mathsf{het}}(15, 116) &= 0.66 \;\; \mathsf{F}_{\mathsf{Chow}}(14, 105) = 0.56 \;\; \mathsf{T} = 1863\text{–}2004.
\end{aligned}$$

The non-linear mapping in (10) is shown in fig. (6). Thus, as price inflation rises, workers become more attentive, and act to prevent the further erosion of their real wages. There is a slow long-run feedback to real unit labor costs, and rather rapid incorporation of productivity increases into real wage increases, dampened by increases in unemployment. The major outliers are all significant. The one counter-intuitive effect is the small but highly significant intercept of about 1% pa ‘autonomous real wage growth’, possibly reflecting unmodeled factors like reductions in hours worked. The fit of 1.24% is only slightly worse than (8), but less close than for (5): over the whole sample, the level of unemployment (and its inverse) were insignificant. Contemporaneous and lagged levels of unemployment and its inverse were included in the GUM, but the reduction procedure found them to be insignificant over the sample

[7] Δtu_p is excluded as its sample only commences in 1870. Some variables only have 1 lag due to sample restrictions, and some contemporaneous values are excluded to avoid simultaneity or perfect collinearity.

[8] Undertaking selection using Autometrics at 1% results in a similar model to equation (11), but the level of unemployment is retained rather than the difference, and a restriction to impose the difference is rejected; the excess demand for unemployment variable is also retained.

period as restrictions testing for differences were all accepted. None of the model mis-specification tests is significant.

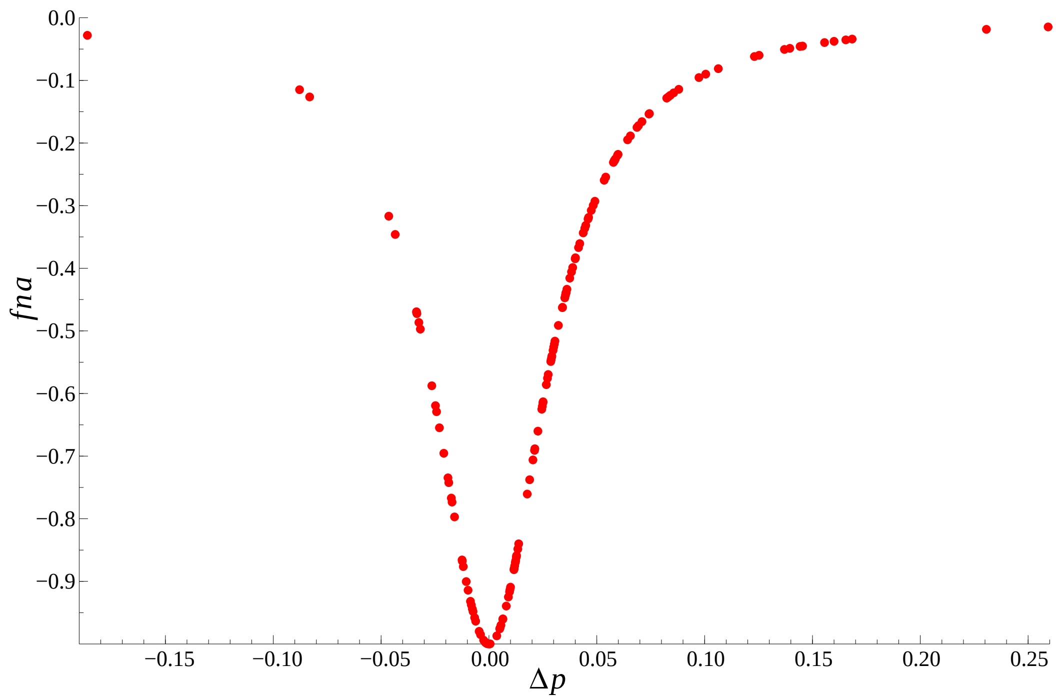

Figure 6 Non-linear price coefficient.

Figure (7) shows the whole-sample outcomes, fit, residuals, their density, and the subsample forecasts. The fit is closest pre-1945 and post the oil crisis, with a deterioration in the 1950s and 60s, although not sufficiently so to induce heteroscedasticity. The 14 'outside sample' 1-step forecasts (i.e., using pre-1990 estimates) are remarkably good given that they are after the Thatcher labor-market reforms, and the 1992 exit from the ERM: despite that large devaluation, and an expectation of higher inflation therefrom that did not ensue, the model does not overpredict.

Finally, fig. 8 shows the resulting recursive estimates and tests. There is evidence of some variation in the estimated parameters beyond pure sampling effects, although the infrequency of 1-step constancy rejections is consistent with not rejecting the null overall. Conversely, extending the 'Phillips model' in (5) leads to $\widehat{\sigma} = 2.0\%$, even with all the dummies added, and a coefficient on Δp_t of unity, so reveals dramatic non-constancy.

7 Conclusions

The breakdown of the original Phillips curve sparked a salient empirical literature on the determinants of wage inflation, of which Sargan (1964) was a seminal paper. Given the performance of these alternative model specifications on different sub-samples over the past 140 years, the question of whether the models can be reconciled is naturally raised. The huge changes that inevitably occur over such a long time period, such as technological innovation, legislative changes, social reforms, wars and policy regime shifts, signify the inherent difficulties in specifing a constant parameter model over the entire 140 years. However, advances in econometric techniques, including the developments in automatic general-to-specific model selection, the use of indicator saturation to identify shifts, and testing for non-linearity in large dimensional, collinear systems, all combine to provide a potential methodology for the analysis

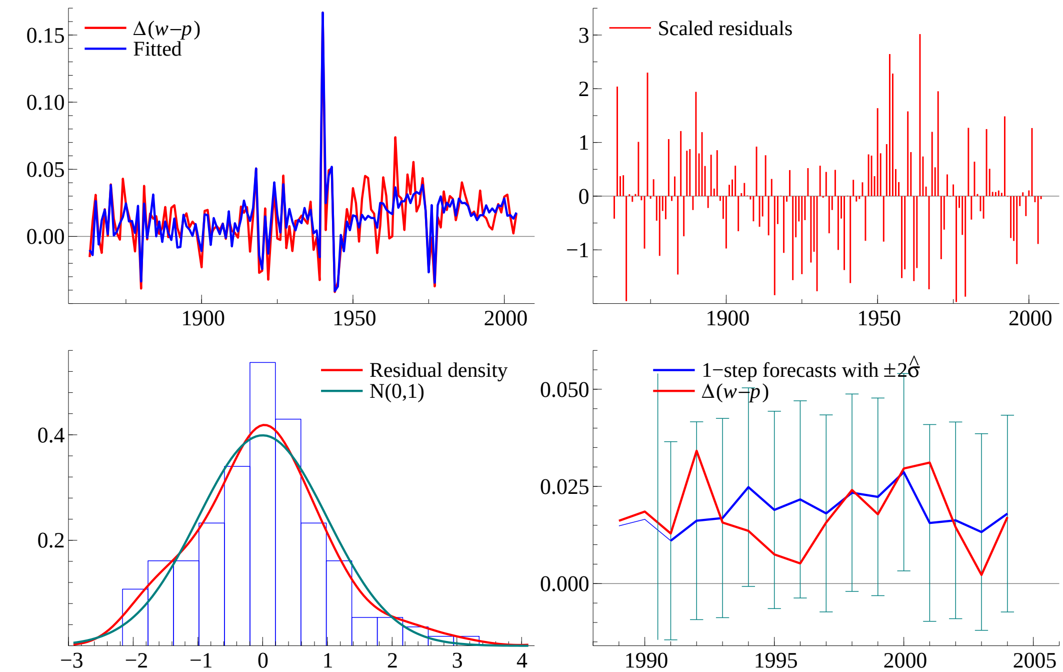

Figure 7 Outcomes, fit, residuals, their density, and subsample forecasts.

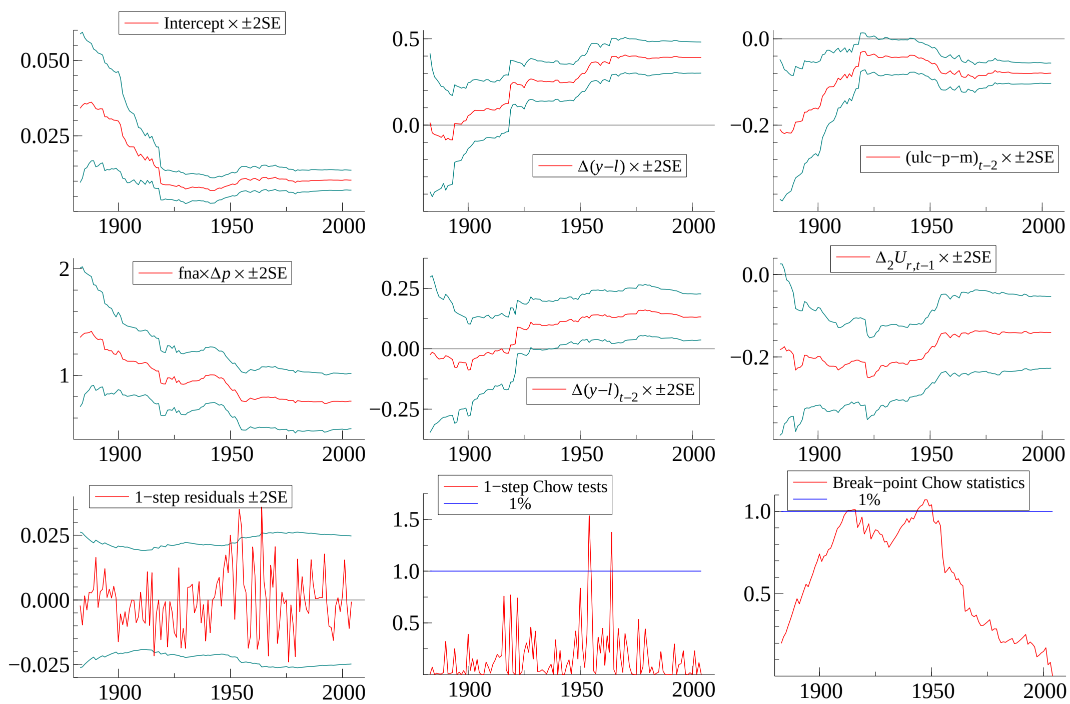

Figure 8 Recursive estimates and tests.

of such a problem. The model of real wage inflation is able to explain much of the movement in wages over almost a century and a half, both pre-World War I and still provide good 1-step forecasts of real wage inflation over the past 14 years. En route, the modeling procedure facilitates tests for many extant theories of wage inflation within a general model. Exclusion of variables such as the replacement ratio and level of unemployment from the final model suggests that these variables do not play a major role in the wage inflation process in the UK. Jacobson, Vredin and Warne (1998) also find little evidence of a relation between real wages and unemployment in post-war Sweden.

Wages are part of an interacting system, some of which has been modelled previously, but all of which has to be understood before policy questions can be successfully addressed. While the results of empirical research always remain tentative, we hope to have thrown some light on the behavior of real wages over the last 144 years.

8 Appendix: Data definitions

Y_t	=	real GDP, £million, 1985 prices	[6], p.836, [9]a (1993)
P_t	=	implicit deflator of GDP, (1985=1)	[6], p.836, [9]a (1993)
M_t	=	nominal broad money, £million	[1], [2]
$R_{s,t}$	=	three-month treasury bill rate, fraction p.a.	[1], [2]
$R_{l,t}$	=	long term bond interest rate, fraction p.a.	[1], [2]
$R_{n,t}$	=	opportunity-cost of money measure	[3]
N_t	=	nominal National Debt, £million	[8]
U_t	=	unemployment	[7], [9]c (1993)
$Wpop_t$	=	working population	[7], [9]c (1993)
$U_{r,t}$	=	$U_t/Wpop_t$ (unemployment rate, fraction)	
L_t	=	employment (= $Wpop_t - U_t$)	[4], [5]
K_t	=	gross capital stock	[6], p.864, [9]c (1972,1979,1988,1992)
W_t	=	average weekly wage earnings	[17], [18], [19]
$W_{r,t}$	=	nominal wage rates	[5], [12], [18]
H_t	=	normal hours (from 1920)	[6], p.148, [9]
ULC_t	=	unit labour costs $(= L_t W_{r,t}/Y_t)$	
TU_t	=	trade union membership (from 1870)	[6], p.137, [15], Table 6.20, [16]
$TU_{p,t}$	=	trade union power $(= 1000 \times TU_t/Wpop_t)$	
B_t	=	unemployment benefits	[13], [14], [12]
$repl_t$	=	replacement ratio $\left(= (b-p)_t \,/\, (w-p)_{t-1}\right)$	
NIC_t	=	national insurance contributions	[12]
$P_{e,t}$	=	world prices, (1985=1)	[1], [2], [10]
E_t	=	annual-average effective exchange rate	[1], [2], [10]
PCL_t	=	cost of living index	[6],p.719,p.737,p.738, [17], [18], [20]
$P_{nni,t}$	=	deflator of net national income, (base 1985)	[1], [2]
$P_{cpi,t}$	=	consumer price index, (1985=1)	[4], [5]
$P_{o\$,t}$	=	commodity price index, $	[11]
U^d_t	=	$U_{r,t} - 0.05 - 0.85\,(R_{l,t} - \Delta p_t - \Delta y_t)$	
y^d_t	=	$y_t - l_t - cap_t$	

$$cap_t = \begin{cases} 2.62 + 0.006t + 0.36\,(k_t - wpop_t) & 1860-1945 \\ 1.69 + 0.015t + 0.36\,(k_t - wpop_t) & 1946-1974 \\ 1.98 + 0.013t + 0.36\,(k_t - wpop_t) & 1975-2005 \end{cases}$$

$$\Delta x_t = (x_t - x_{t-1}) \text{ for any variable } x_t$$

$$\Delta^2 x_t = \Delta x_t - \Delta x_{t-1}$$

Sources: [1] Friedman and Schwartz (1982); [2] Attfield, Demery and Duck (1995); [3] Ericsson *et al.* (1998); [4] Shadman-Mehta (1995) (who cites Sleeman (1981) and Thomas (1984) as sources); [5] Phillips (1958); [6] Mitchell (1988); [7] Feinstein (1972); [8] Bank of England; [9] Bean (taken from (a) *Economic Trends Annual Supplements*, (b) *Annual Abstract of Statistics*, (c) *Department of Employment Gazette* and (d) *National Income and Expenditure*, as well as other sources cited here); [10] Cameron and Muellbauer; [11] UN Statistical Yearbook; [12] Office for National Statistics, Blue Book; [13] Board of Trade (1860-1908); [14] SS Stats; [15] Annual Abstract of Statistics; [16] Office for National

Statistics, Labour Market Trends; [17] Crafts and Mills (1994); [18] Feinstein (1990); [19] Office for National Statistics, Labour Force Survey; [20] Office for National Statistics, Economic Trends Annual Supplement.

Notes:

Hendry and Ericsson (1991) and Hendry (2001) provide detailed discussions about most of these series.

Average weekly wages: a measure of full-time weekly earnings for all blue collar workers, where the coverage has been extended to include more occupations. and allows for factors such as changes in the composition of the manual labour force by age, sex, and skill, and the effect of variations in remuneration under piece rates and other systems of payments, but not adjusted for time lost through part-time work, short-time, unemployment etc. A reduction in standard hours worked that was offset by a rise in hourly wage rates would not be reflected in the index. From 1855-1880, the data are from Feinstein (1990), but not revised to increase coverage. Prior to that, the data come from a number of sources on average wage rates for blue collar workers.

Nominal wage rates: hourly wage rates prior to 1946, then weekly wage rates afterwards, so the latter were standardized by dividing by normal hours. The trend rate of decline of hours is about 0.5% p.a. (based on a drop from 56 to 40 between 1913 and 1990, with an additional increase in paid holidays), so unit labour costs were adjusted accordingly, and spliced to an average earnings index for the whole economy including bonuses [ONS: LNMM] from 1991 and rebased to 2000=1.

Benefits: amount expended for poor relief in unionized industries/number of paupers (from [13]) to 1908 (data for England and Scotland). Interwar and post-WWII data on value of benefits from [14], Table 34.01 (1989), Table A2.36 (1992) and [12] total government benefit expenditure/population. Data spliced over 1909-1919 and 1939-1945.

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