Do minimum wages raise the NAIRU ?

Peter Tulip[1]

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Abstract

Probably yes.

A high minimum wage (relative to average wages) raises nominal wage growth and hence inflation. This effect can be offset by extra unemployment; so the minimum wage increases the Non-Accelerating Inflation Rate of Unemployment or NAIRU.

This effect is clearly discernible and robust to variations in model specification and sample period. It is consistent with international comparisons and the behavior of prices.

I estimate that the reduction in the relative level of the minimum wage over the last two decades accounts for a reduction in the NAIRU of about 1½ percentage points. It can also account for the substantial reduction in the NAIRU in the USA relative to continental Europe.

[1] Division of Research and Statistics, Federal Reserve Board, Washington D.C. 20551. Email: ptulip@frb.gov. I am grateful to Flint Brayton, Dave Reifschneider, Albert Ando, Lawrence Klein, Dan Polsky, Charles Fleischman, Frank Schorfheide and seminar participants at the Federal Reserve Board for helpful comments and assistance. This paper is a revised version of Chapter 1 of my PhD Dissertation. Related chapters and my data are available at http://petertulip.homepage.com. The views in this paper do not necessarily represent those of the Federal Reserve Board or its staff.

1 INTRODUCTION

Estimates of the Non-Accelerating Inflation Rate of Unemployment, or NAIRU, serve several purposes. Central bankers such as Gramlich (1998) and Blinder (1997) have used them to guide monetary policy. Stiglitz (1997) argues that they are useful for forecasting inflation and framing policy discussions. The OECD (1994) and Layard, Nickell and Jackman (1991) use them as a measure of the permanent component of unemployment.

In this paper, I investigate whether variations in the NAIRU can be explained by variations in the level of the minimum wage (relative to the average wage). I find that they probably can. I find the effect of the minimum wage to be important, clearly discernible and robust. On average, a 10 per cent increase in the minimum wage, relative to average wages, seems to raise the NAIRU about half a percentage point. Movements in the level of the minimum wage help explain the upward drift of the NAIRU in the USA over the 1960s and 1970s and its subsequent decline. They also help explain why the NAIRU has risen in continental Europe while falling in the USA. Much of the surprisingly favorable behavior of inflation and unemployment in the USA over the 1990s can be attributed to the low level of the minimum wage.

Although the effect of the minimum wage on the NAIRU is relevant to macroeconomists, it may be of greater interest to those assessing minimum wage policy. Specifically, it measures the sustainable unemployment arising from the minimum wage. Because accelerating prices cannot be sustained, the long-run level of unemployment arising from the minimum wage will approximately equal its effect on the Non-Accelerating Inflation Rate of Unemployment.

This measure of the consequences of the minimum wage may be more useful than the conventional focus on changes in employment. (For surveys of the conventional approach see Brown, Gilroy and Kohen, 1982 or Card and Krueger, 1995, Chapters 6 and 7). One limitation of job loss estimates is that they focus upon “affected workers” and say little about overall effects. A more fundamental limitation is that they ignore Phillips curve effects. Typically, increases in unemployment put downward pressure on wages. If unemployment arising from the minimum wage is similar to unemployment arising from other influences, it would imply continually falling wages. This would increase employment elsewhere and the unemployment would be temporary. So, unless the minimum wage somehow shifts the Phillips curve, estimates of immediate job losses will overstate the total effect of the minimum wage. Of course, such a shift is possible – indeed, the subject of this paper. But this can be estimated directly. The short-term change in employment of affected workers provides little information about the long-term change in total unemployment, which seems to be the variable of most interest to policy makers.

I am aware of two previous attempts to estimate the effect of the level of the minimum wage on the NAIRU. Jackman, Layard and Nickell (1996 n2) state "It would be very desirable to find a way of including minimum wages [in their NAIRU-like framework] but we have not found a satisfactory way to do so." And Staiger, Stock and Watson (1996, p26) find that the nominal minimum wage is insignificant in explaining the acceleration of prices in the USA, but do not report estimates. My research differs from these attempts in that I find a significant effect of the ratio of the minimum wage to average wages (adjusted for coverage) in an equation explaining the growth in nominal wages. If this effect flows on to prices – as it appears to do – then extra unemployment would be required to offset the inflationary pressure, so the NAIRU is higher.

Other researchers have uncovered similar empirical relationships. Adams (1989) finds that the gap between average and minimum wages is significant in a wage equation. And an apparent "tradeoff" between inequality and unemployment has been noticed in many cross-country comparisons (section 3a gives citations). However, previous researchers have not interpreted these relationships as an effect of the minimum wage on the NAIRU, nor have they examined their robustness. Less directly related, Gramlich (1976), Grossman (1983), Spriggs and Klein (1994), Card and Krueger (1995) and many others have found "ripple" effects from the minimum wage to other wages. However these other studies have not looked for (or at least, do not report) the persistent, and hence larger, effects that I detect.

Merely presenting one partial correlation would be unconvincing, given the skepticism with which new econometric results are received. Leamer (1983), McAleer, Pagan and Volker (1985), Hendry and Mizon (1990) and others have discussed how the credibility of econometrics might be enhanced. Proposals include sensitivity analysis, examining various kinds of evidence and diagnostic testing. I follow some of this advice in this paper, the plan of which is as follows. Section 2 defines my measure of the NAIRU, explains how estimates of it can be derived from a wage equation and presents those estimates. Section 3 shows that my results are consistent with other sources of information. Specifically, the effect of the minimum wage on the NAIRU is also evident in the behavior of prices and in international comparisons. Section 4 looks at the estimates from the wage equation in detail and shows that they are robust to variations in model specification, sample period and policy. Section 5 concludes.

As the plan indicates, the focus of this paper is empirical. It seems useful to determine whether minimum wages have an important effect on the NAIRU before asking how or why this might be so. Nevertheless, it may be interesting to note a few reasons why such an effect might be expected.

The minimum wage might increase the NAIRU by reducing the demand for unskilled workers or by encouraging potential workers to look for work. If these newly unemployed workers are ineffective searchers – perhaps because they are poor substitutes for other workers – they might put little downward pressure on wages. Higher unemployment would be associated with a similar wage outcome.

The minimum wage can also increase the NAIRU if it causes other nominal wages to gradually increase, at a given rate of unemployment, and if these wage increases are passed on to higher prices. The literature on "ripple" effects has attributed them to considerations of fairness and a desire to recruit, retain or motivate workers. For formalizations, see Grossman (1983) or Bulow and Summers (1986). There is also support for less obvious channels of influence. For example, Card and Krueger (1995 p163) suggest that the minimum wage might determine starting wages but not the shape of wage-tenure profiles. Then an increase in the minimum wage would slowly flow to higher aggregate wages as workers gain promotions. Another possibility is that the minimum wage acts as a safety net or "outside option" that affects workers' bargaining position. This interpretation seems consistent with (and would help to explain) the widespread view that the reduction in the NAIRU in the 1990s was due to worker insecurity.

In Tulip (2000, Ch3) I discuss these and some other potential interpretations. I note that a difficulty with attributing large effects of the minimum wage to the demand or supply of labor is that relatively few workers are paid the minimum wage. However this objection does not apply to effects based on interaction of wages, as a shock initially affecting a small group of workers can flow on to others. But these matters of interpretation lie outside the scope of this paper.

2 FRAMEWORK AND ESTIMATES

2a. Definition of the NAIRU

The NAIRU is the unemployment rate consistent with stable inflation. Loosely speaking, it is estimated as the horizontal intercept of a Phillips-curve. Any variable that shifts out the Phillips curve increases the NAIRU. More unemployment is required to offset the inflationary impact of the shock. Thus to find out whether the minimum wage raises the NAIRU essentially involves testing whether minimum wages are a significant determinant of inflation.

Most estimates of the NAIRU are based on equations in which inflation is regressed on lagged inflation and other variables, often called a “price-price Phillips curve”. As I show in section 3b, my results can be obtained from an equation in this style. Price-price Phillips curves tell a simple story and can be useful for forecasting. However, because they omit apparently important influences, such as wages, they are difficult to interpret. For policy analysis, my preference is to explain prices as a markup on unit labor costs and then focus on the determinants of nominal wage growth. This approach means one does not have to control for product market influences that are difficult to measure. It also provides more easily interpretable estimates and better explains some of the data. It does however require assuming that increases in labor costs are fully passed on as increases in prices.

I outline below a simple model in which estimates of the determinants of the NAIRU can be obtained from interacting price and wage equations. This three-equation system resembles the price-wage block of old macroeconometric models, such as the MPS model. An earlier version was presented in Ando and Brayton (1995).

The main equation explains the growth of nominal wages, W, in terms of past growth in consumer prices, P_c, the unemployment rate, U, the growth of trend productivity, *prod*, and a vector X of other factors including a constant, the minimum wage and residual. Measuring these variables in logarithms, letting Δ represent the first difference operator and other Greek letters represent elasticities, then the wage equation can be written:

$$\Delta W = \Delta P_{c(-1)} + \alpha \Delta prod + \beta U + \delta X \qquad (1)$$

This equation is intended to capture the main empirical influences governing wage growth – though how well it does so clearly depends on how X is constituted. I discuss its specification in the following section. The equation is not derived from a clearly defined optimization problem. The preferences and constraints that give rise to the inertia evident in the wage data remains a subject of active research.

The wage equation is important because, in the medium to longer run, prices appear to mimic wage movements. As I discuss in Appendix 1, the behavior of log product prices, P_p, in the modern U.S. economy, can be described well by an equation of the form:

$$\Delta P_p = \Delta(W - prod) + \lambda(P_p - W + prod)_{(-1)} + \gamma U + Z \qquad (2)$$

where $(W - prod)$ represents the logarithm of nominal unit labor costs and Z represents the effect of the Nixon price controls, changes in energy, farm and import prices and a small, stable residual. To simplify the exposition I ignore most lags and constrain the coefficient on the change in unit labor costs to unity. The critical feature of this equation is the role played by the *levels* of prices and wages. An “error-correction mechanism” adjusts prices so as to gradually eliminate deviations from their steady state level, P_p*. The steady state markup is given as:

$$P_p^* = W - prod - (\gamma U + Z) / \lambda \qquad (3)$$

Empirical estimates of equation (2) are presented in Appendix 1. These imply that, on average, 12 per cent of any deviation of prices from their target P_p* is closed each quarter. So, in the long run, P_p will approximately equal P_p*. In the modern U.S. economy, U and Z are close to stationary (as a long-run approximation $\Delta U = \Delta Z = 0$), as are the inflation rate, the relative minimum wage and many other variables that could be included in the specification. Using these conditions and differencing (3) gives a long-run version of equation (2):

$$\Delta P_p = \Delta W - \Delta\, prod \tag{4}$$

The third element of the model is an equation linking product prices to consumer prices:

$$\Delta P_c = \Delta P_p + \; wedge \tag{5}$$

The difference between consumer prices and product prices, which I call *wedge*, is a composite of numerous influences that I take as exogenous. In the short term, the most important of these include fluctuations in farm prices and the external terms of trade. In the longer term, differences in productivity trends between different sectors are important. In particular, faster technological change in the production of investment goods relative to consumption goods has meant that the wedge has been positive since the early 1980s. (It was near zero, on average, before this).

Substituting (1) into (2), then (2) into (5) would give a reduced form for inflation in terms of lagged inflation – from which a NAIRU could be estimated. However, imposing the long-run empirical regularity (4) gives an equation that is simpler and more easily interpreted. Substituting (1) into (4), then (4) into (5) gives a reduced form that applies once price margins have returned to their long-run levels:

$$\Delta P_c = \Delta P_{c(-1)} \; + \; (\alpha\text{-}1)\Delta prod + \beta U + \delta X + \; wedge \tag{6}$$

The NAIRU is defined as the unemployment rate at which inflation is stable. Setting $\Delta P_c = \Delta P_{c(-1)}$ and solving for unemployment gives:

$$NAIRU = \; -[\delta X + wedge + (\alpha\text{-}1)\Delta prod\,] \,/\; \beta \tag{7}$$

This definition of the NAIRU is essentially determined in the labor market. It is independent of many product market shocks (specifically, the Z variables) and the effect of unemployment on price margins (the parameter γ). The reason is that these factors cause product prices to grow at a different rate to unit labor costs. This causes a deviation of price margins from the long-run level given by equation (3), which is offset by the error-correction mechanism.

Equation (7) provides a framework for estimating contributions to the NAIRU. Specifically, I decompose the vector X into those variables that seem to be important for wage determination; I estimate the parameters δ, α and β from an equation like *(1)* and I take a long-term approximation of *wedge*. I can ignore short-term influences on product and consumer prices.

This framework represents a tradeoff between realism and simplicity. A simpler model, for example with $\alpha = 1$ or *wedge* = 0, would have weaker explanatory power. Tempting generalizations include the addition of vacancies and the lagged wage share to the wage equation and allowing for non-stationary elements in Z. These complications seem to be important in other OECD economies, though not in the USA.

Suppose the minimum wage is an element of X with coefficient δ^*. Then the fundamental parameter of interest is δ^*/β – the contribution of the minimum wage to the NAIRU. However, given that I have strong priors about the coefficient β (a small negative value is uncontroversial), I am especially interested in estimates of δ^* – the contribution of the minimum wage to nominal wage growth. Much of the framework outlined above is not necessary to determine this. Nevertheless, other estimates (for example, of the level of the NAIRU) provide some interesting context and help in comparing my results to other research and data.

2b. Wage equation estimates

An econometric equation explaining nominal wage growth is set out below. This is a modified version of the wage equation of the old MPS model, as published in Ando and Brayton (1995). The dependent variable is the quarterly change in the logarithm of average hourly compensation in the non-farm business or private sector in the USA from 1948 to 1998. Explanatory variables include productivity growth, inflation, unemployment, the relative minimum wage, unemployment benefits, payroll taxes and the 1971 wage freeze. The coefficients on inflation are constrained to sum to 1. An interaction term raises the effect of the change in the minimum wage when the level of the minimum wage is high.

My measure of wages splices data from the BLS's Productivity and Cost release and the Employment Cost Index at 1980. This gives lower standard errors for the wage equation and, more importantly, the price equation (Appendix 1) than relying on the Productivity and Cost measure alone. This suggests that the spliced series has less measurement error, permitting more precise estimates and more robust inferences. It also exacerbates the heteroskedasticity in the data, for which I control by weighted least squares. Weights are jointly estimated by maximizing a likelihood, assuming that the residuals are normally distributed with a variance that declines exponentially over time. The procedure is described in Harvey (1976). My estimates are as follows (with standard errors in square brackets under the respective coefficient):

$$
\begin{aligned}
\Delta\text{Wage} = \;& \underset{[.0016]}{.0084} \log(\text{minimum wage x coverage})_{(-1)} \\
+\;& (\underset{[.12]}{.65} + \underset{[.037]}{.19} \log \text{minimum wage}_{(-2)}) \text{ x coverage}_{(-2)} \text{ x change in minimum wage} \\
+\;& \underset{[.0011]}{.0018} \text{ unemployment benefit replacement rate}_{(-1)} \\
+\;& \underset{[.20]}{.53} \text{ trend productivity growth} \\
+\;& \underset{[.05]}{.53} \text{ last year's inflation} \\
+\;& (1 - \underset{[.05]}{.53}) \text{ previous four years' inflation} \\
-\;& \underset{[.00054]}{.0023} \text{ change in unemployment rate}_{(-1)} \\
-\;& \underset{[.00022]}{.0014} \text{ unemployment rate}_{(-2)} \\
+\;& \underset{[.20]}{.88} \text{ change in employer's social security contribution rate} \\
-\;& \underset{[.002]}{.012} \text{ dummy for 1971 wage freeze} \\
+\;& \underset{[.069]}{.031} \text{ (constant)}
\end{aligned}
$$

The residuals are assumed to be distributed normally, with variance given by:

$$\log(\sigma^2) = -10.36 - .0155\ \text{TIME}.$$
$$\quad\quad\quad\quad [0.19] \quad\quad [.0017]$$

Summary statistics and diagnostics: [2]

sample	1948:Q3 1998:Q2
standard error	0.002435
R-Squared	.88
RESET test of functional form	1%
Breusch-Godfrey test for 4th order serial correlation	17%
White's test for heteroskedasticity/misspecification	9%
Jarque-Bera test for normality	56%
Andrews-Ploberger test for stability (1951:Q3 - 1995:Q3)	13.7 (p <1%)
Andrews-Ploberger test for stability (1963:Q3 - 1995:Q3)	7.6 (p >10%)

Appendix 2 provides details of data measurement and sources. In brief, the minimum wage is the statutory minimum divided by average hourly compensation. Coverage is the proportion of employees covered by the Federal minimum wage legislation. Productivity growth is a five year back-average of the quarterly change in trend log business output per hour, with a kink at 1973:Q1. "Last year's inflation" is the average change in the logarithm of the consumption deflator over the previous four quarters. The unemployment rate is in percentage terms, with fixed demographic weights.

All coefficients have p-values less than 1% (assuming normality) except the unemployment benefit replacement rate, which has a p-value of 10%. Diagnostics are reported as p-values of F-statistics (interpretable as the probability that the sample is consistent with the null hypothesis of no problems), except for the Andrews-Ploberger

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[2] These are based on weighted residuals (that is, divided by the estimated standard error). They are calculated by Eviews 3.1, with the following exceptions. RESET is an F-test of the addition of the squared fitted values to the equation. Breusch-Godfrey is an F-test of the addition of four lagged residuals to the equation, with presample values of lagged residuals set to zero. Both tests use the same weights as the primary equation. Andrews-Ploberger is the Exp-W_∞ statistic discussed in Andrews, Lee and Ploberger (1996, equation 2.15) – essentially a weighted average of Wald breakpoint tests – with coefficients on the wage freeze, productivity and the standard error trend constrained to their full-sample estimates. I present this statistic both for all feasible breakpoints and for a large range over which it is insignificant. Critical values, as tabulated by Andrews and Ploberger (1994) for their parameters of p = 9 and λ = 7, are 11.4 (1%) and 8.1 (10%).

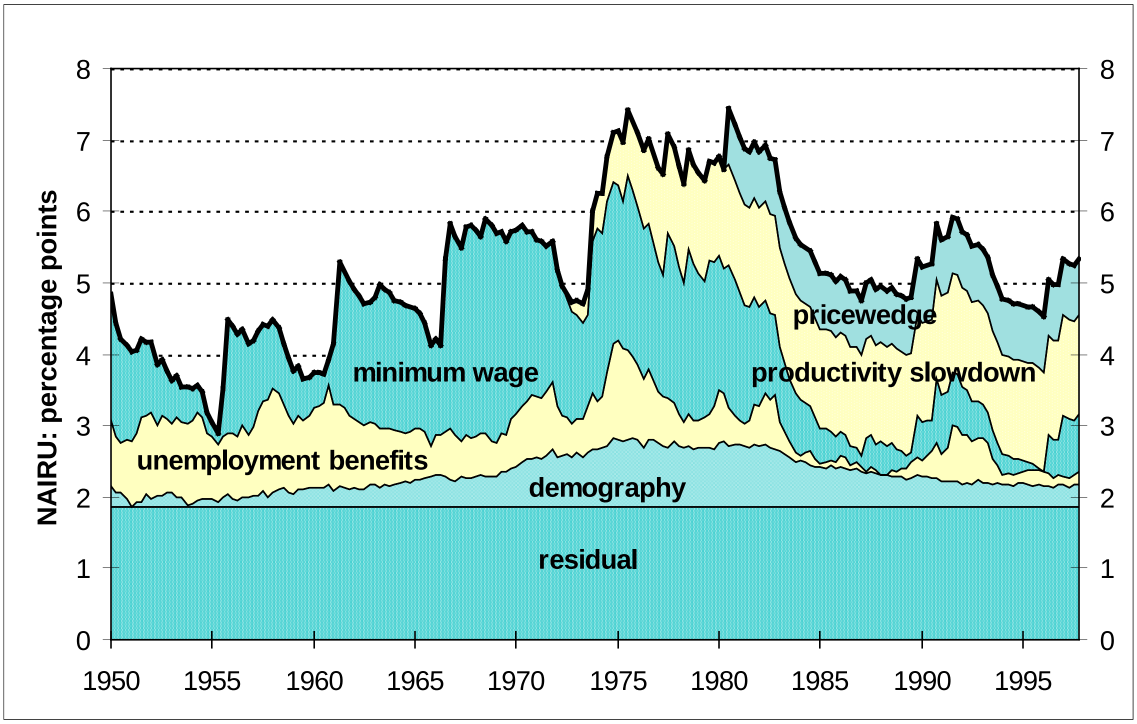
test of parameter stability, which has a non-standard distribution. I discuss tests of instability in section 4a and mis-specification in section 4b.

2c. NAIRU estimates

The definition of the NAIRU in equation (7) could be interpreted as including temporary influences on wages; specifically changes in unemployment, payroll taxes and the minimum wage, incomes policy, the residual and the different lags of adjustment of wages and prices to the break in the productivity trend. These variables can have large impacts on the unemployment required to stabilize labor costs for short periods of time. For example, to offset a percentage point increase in the social security contribution rate requires an extra percentage point of unemployment be maintained for 6 (=.88/.0014 x 1%) quarters. However, because such influences are frequently reversed, they have a small total effect. Averaged over the full sample period, they raise the NAIRU by half a percentage point.

Because transitory influences obscure the mean of the NAIRU and because a shock that is soon reversed does not need to be offset by a change in unemployment, it seems more interesting to focus on persistent contributions. This is consistent with the treatment of prices and facilitates comparison with other research. Accordingly, Chart 1 below shows estimated contributions to the NAIRU from the constant, the level of the minimum wage, unemployment benefits, demography, productivity and the trend price wedge since 1950. The sum of these contributions equals the rate of unemployment consistent with stable inflation in the long run. By “the long run” I mean a time horizon long enough for deviations in price margins and temporary influences on wages to be unimportant. This might be between a few years to a decade.

Chart 1: US NAIRU – Persistent contributions



Contributions in the chart are measured as deviations from the minimum value observed since 1950. This is a graphical device to ensure each series is positive; accordingly, the level of each contribution, at any point in time, is of little interest. Rather, the interesting feature is how these contributions change over time. In particular, the chart shows the growing contribution of the minimum wage over the 1960s and 70s and its subsequent decline. I discuss the size of minimum wage effects in the next section.

The formula for the NAIRU is obtained by substituting estimated coefficients from the wage equation into the definition of the NAIRU in equation (7):

$$\text{NAIRU} = \underset{[0.6]}{6.3} + \underset{[1.3]}{5.9}\text{ minimum wage} + \underset{[0.7]}{1.3}\text{ UI benefits} - \underset{[186]}{328}\text{ productivity} \quad (8)$$

$$+ \underset{[0.002]}{\text{wedge}/0.0014} + \text{demography}$$

The wedge is zero to 1981 then 0.0011 thereafter. Demography is the difference between the unemployment rate and a rate with fixed demographic weights. The minimum wage and unemployment insurance benefits are here measured as deviations from their sample mean (in contrast to the chart), to facilitate interpretation of the constant. Both variables are in logarithms; so, for example, a 10 per cent increase in the minimum wage increases the NAIRU by approximately 0.59 percentage points.

Standard errors, calculated by the delta method, are in square brackets. Because the coefficients are ratios of correlated least squares coefficients they will not be normally distributed. I estimate a 95% confidence interval for the NAIRU at the end of my sample period, 1998:Q2, extends from 4.8 to 5.6 per cent.[3] This is narrow relative to the standard error of the constant because of negative covariances between coefficients. It is also narrow relative to the confidence intervals reported by Staiger, Stock and Watson (1996). This seems to reflect the use of wage rather than price data, and the long lags in my specification, as discussed in Brayton, Roberts and Williams (1999). Further measures of uncertainty could be calculated – for example, to allow for unmodeled heteroskedasticity or serial correlation – however these issues seem to be unimportant relative to uncertainty about whether the model is correctly specified. I discuss specification uncertainty in section 4.

Estimates of the level of the NAIRU are not my primary interest. Nevertheless, it is reassuring that my estimates conform to those of others who have looked at this more closely, including Tobin (1980 p58), Stiglitz (1997, p6), Gordon (1997), Staiger, Stock and Watson (1996), the FRB/US model (1999), the OECD (1996b, Table 1) and Laubach (forthcoming). There are many differences between (and within) these studies; however there is some agreement that the NAIRU increased over the 1960s and 1970s, peaked near 6½ per cent around 1980, then declined by 1 to 2 percentage points to the mid 1990s.[4] These results corroborate the main trends in Chart 1. Furthermore, my

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[3] I estimate this by taking 10,000 draws of a random vector that has a multivariate normal distribution with mean and covariance equal to the coefficients in my wage equation. I substitute these into (7) to obtain 10,000 realizations of the coefficients in equation (8). Taking the 1998:Q2 values of the variables in equation (8) as given, I then calculate 10,000 realizations of the NAIRU for this period.

[4] The main disagreement concerns the late 1950s, when Gordon estimates the NAIRU was around 6 per cent and falling. In contrast, Tobin puts the NAIRU at 3 per cent in the early 50s rising to 4 per cent in the 1960s.

decomposition provides an explanation of other researchers' results and hence provides a basis for analysis and forecasting.

2d. The size of the minimum wage effect

Were an increase in the relative minimum wage of 0.1 log points (approximately 10 percent) enacted in mid-1998, I estimate it would have immediately raised aggregate nominal wages by 0.3 per cent. Then, the higher level of the minimum wage would continue to raise aggregate wages by 0.084 per cent a quarter.

This continued effect seems small. It is less than rounding errors in the data and substantially less than the standard error of my wage equation. However, it is substantial relative to the effect of unemployment on wages. To offset a 0.1 log point increase in the relative minimum wage requires an extra 0.59 percentage points (0.0084/.0014 x 10%) of unemployment. Thus the 0.3 log point reduction in the coverage-adjusted relative level of the minimum over the 1980s accounts for a reduction in the NAIRU of about 2 percentage points. Comparing the late 1970s with the late 1990s, the lower minimum wage reduced the NAIRU by about 1½ percentage points. Congress is currently (as of early 2000) considering an increase in the Federal minimum wage from $5.15 to $6.15 an hour. I estimate that this would raise the NAIRU by one percentage point, relative to a policy of no change in the nominal minimum.

The size of these estimates has surprised many American readers. This may be because the flatness of the short-run Phillips curve is sometimes not appreciated. It actually takes a lot of unemployment to offset small supply shocks. It may also reflect the novelty of the results. Until recently, the NAIRU in the USA was thought to be constant and researchers have had difficulty in detecting effects of minimum wage changes. In contrast, economists outside the USA are accustomed to large variations in the NAIRU; these are commonly attributed to institutional factors.

Although it may seem small, the effect of the level of the minimum wage on nominal wage growth is precisely estimated and clearly discernible. The coefficient on the level of the minimum wage in the wage equation is over 5 times as large as its estimated standard error. If the coefficient is normally distributed[5] its p-value would be less than 0.001 per cent. The estimated effect of the minimum wage on the NAIRU is non-normal and slightly less precise. I estimate a 95% confidence interval for this coefficient spans 3.6 to 9.0.

3 CORROBORATING EVIDENCE

3a. International comparisons

In many other countries governments strongly influence the wages of low-paid workers (though typically not through national statutory minima, as in the USA). Their experience provides a means of corroborating the patterns evident in U.S. data. One way of quantifying and comparing the effect of this intervention is through the ratio of the 10th percentile of the wage distribution to the median. In U.S. data, this "10/50 ratio" is strongly correlated with the ratio of the minimum wage to the mean – as would be expected given the proximity of the minimum and mean wage to the 10th and 50th percentiles respectively. The OECD's 1996 Employment Outlook (Table 3.1) provides annual estimates of the 10/50 ratio from around 1980 to the mid 1990s[6]. Unfortunately, this period comes after some of the largest compressions of wage relativities were enacted (for example, in France, the Netherlands, the UK and Australia).

I compare these changes in relative wages with time-varying estimates of the NAIRU compiled by the OECD Secretariat (1996b, Table 1). These are based on the bivariate relationship between inflation and unemployment, adjusted to reflect additional research, where available. Earlier estimates from this series have been published and discussed by Elmeskov (1993) and Ball (1996). Laubach (forthcoming, Figure 5) presents similar estimates for seven of these countries, using a Kalman filter and controlling for changes in exchange rates and commodity prices.

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5 Although the regressor is the ratio of two integrated variables, the work of Kremers, Ericsson and Dolado (1992) suggests this assumption is reasonable.

6 Data relate to both sexes, except for the United States, where I use the average of estimates for men and women. Where estimates for 1980 or 1995 are not available, I take nearby estimates.

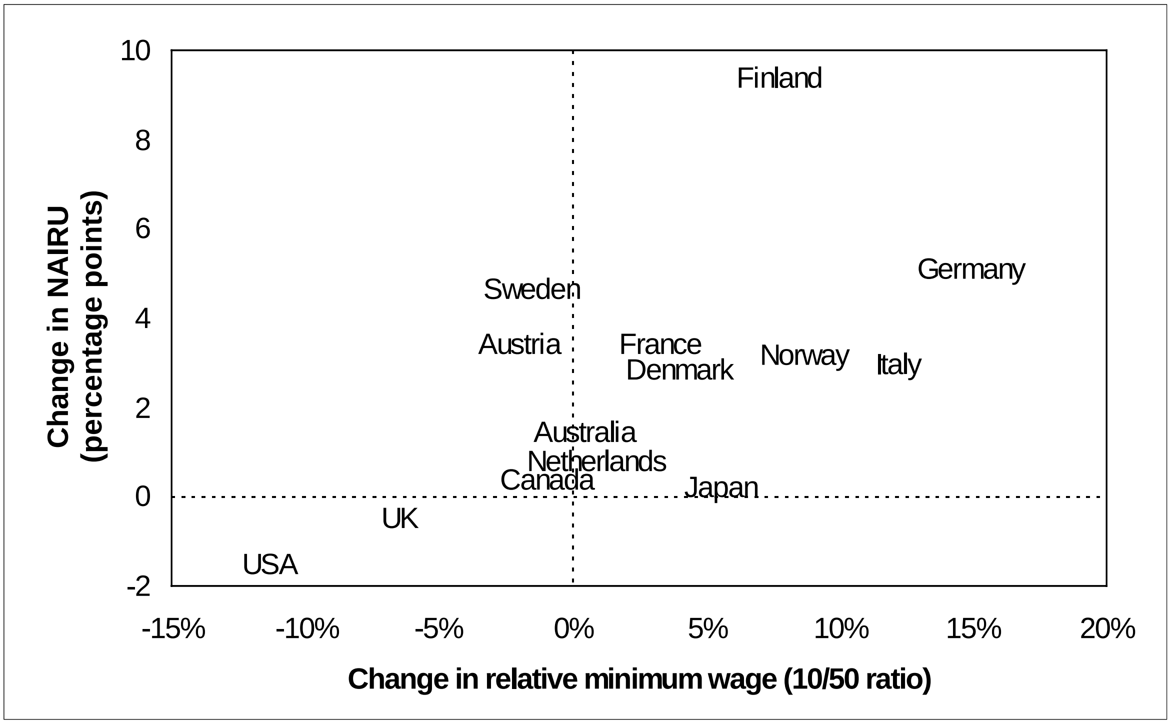
To be convincing, the effect of minimum wages on the NAIRU should be embedded in a multivariate framework. That is an advantage of a close focus on one country's data. However, models in which this could be implemented have not yet been developed. Nor has the literature on the determinants of the NAIRU indicated influences that clearly need to be controlled for. Given this, I begin by examining the bivariate relationship.

Siebert (1997 p51) claims that countries with explicit economy-wide minimum wages have high unemployment rates. However, I find no clear relationship between levels of inequality and levels of the NAIRU for particular points in time, despite wide variations in my data set. This may be because of the inconsistent measurement of wage relativities across countries, because other factors overwhelm the relationship and/or because the relationship is actually unimportant.

Comparisons over time do not suffer as severely from the problems of inconsistent measurement or fixed effects and so are more informative. Chart 2 below plots changes in the NAIRU against changes in the 10/50 ratio between 1980 and 1995. Countries where wages at the bottom of the distribution have fallen relative to other wages – that is, the USA and the UK – have experienced reductions in the NAIRU. Meanwhile, countries where wages at the bottom of the distribution have risen, as in Europe, have seen NAIRU increases.

Chart 2: Changes in relative wages and the NAIRU

OECD Economies; 1980 to 1995



The international comparisons convey the same message as the U.S. wage data: when wages at the bottom of the distribution are compressed, the NAIRU usually rises. The relationship is also similar in quantitative terms. Chart 1 indicated that if wages at the bottom of the distribution had kept pace with wages at the middle after 1980 then the U.S. NAIRU would have been about 2 percentage points higher in 1995. The experience of other countries, presented in the chart above, suggests a similar effect.

The noise in the bivariate relationship, and the apparently positive vertical intercept, indicate that influences on the NAIRU other than wage relativities are important. (The OECD-wide productivity slowdown may explain the positive intercept). Nevertheless, the large variations in the NAIRU allay concern about omitted variable bias. Other variables would need to have very large effects in order to explain away this correlation.

Many observers have noticed variations on the relationship shown in Chart 2. The OECD Jobs Study (1994, Chart 5.1), for example, shows a similar relationship between changes in the 90/10 ratio and the growth in private sector employment. Other descriptions of the apparent "tradeoff" between jobs and inequality include Krugman (1994), Freeman (1995), Bertola and Ichino (1995), Mortensen and Pissarides (1997) and Blank (1997).

Several suggested explanations of this relationship appeal to institutional differences to explain how a common exogenous shock has resulted in wider wage dispersion in the United States and Britain, but higher unemployment in Europe. However, attempts to identify this external shock have often been judged to be unsuccessful (see, for example, the discussion following Bertola and Ichino (1995)). A simpler explanation would be that the relationship between unemployment and inequality is causal. Part of the increase in unemployment in Europe may be due to increases in the relative wages of low-paid workers.

3b. A "price-price" Phillips curve

Most estimation of the NAIRU in the USA is conducted in the framework of a "price-price" Phillips curve in which inflation is regressed on its own lagged values. Such an equation could be estimated jointly with my wage equation to obtain more efficient estimates. However, I am more interested in inference than in estimation, and I use a price-price equation for cross-validation.

Estimates of a price-price Phillips curve are presented below – in tabular form, which is more convenient for a linear specification. The specification is similar to models published by Gordon (1997), Fuhrer (1995), Wiener (1993), Akerlof, Dickens and Perry (1996) and Brayton, Roberts and Williams (1999); the main innovation is the inclusion of the level of the minimum wage. The dependent variable is the quarterly log difference in the chain-weighted price index for personal consumption expenditures.

Table 1: A Price-Price Equation[7]

Regressor	Coefficient	Standard error
last year's inflation	.63	.044
previous five years' inflation	.27	
relative level of minimum wage(-1)	.0031	.0012
unemployment rate(-1)	-.00061	.00010
food and energy prices	.77	.063
change in relative import prices	.040	.011
dummy for Nixon price controls	-.0011	.00031
constant	.014	.0043

sample	1952:Q3 1998:Q2
standard error	0.0020
R-Squared	.91
RESET test (with squared fitted values)	66%
Breusch-Godfrey test for 4th order serial correlation	3%
White's test for heteroskedasticity/misspecification	99%
Jarque-Bera test for normality	61%
Andrews-Ploberger test for stability (1954:Q2 - 1996:Q4)	11.5 ($p < 1\%$)
Andrews-Ploberger test for stability (1958:Q2 - 1996:Q4)	4.6 ($p > 10\%$)

The level of the minimum wage is again statistically significant, with a p-value of 1 per cent (assuming normality). As with unemployment, it has a slower effect on prices than on wages, presumably reflecting the less direct channel of influence. Dividing the coefficient on the minimum wage by that on unemployment gives a minimum wage coefficient in an equation defining the NAIRU of 5.1 (standard error 2.2). This is statistically and economically similar to the estimate from the wage equation (5.9, standard error 1.3).

It is possible to estimate a price-price equation in which the effect of the minimum wage is less clear. Perhaps the strongest alternative specification in which this is the case

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7 Statistics are the same as for the wage equation but without adjustments for heteroskedasticity. The Andrews-Ploberger tests are based on Chow statistics and Andrews-Lee-Ploberger's (1996, equation 4.3) Exp-F_∞ statistic.

includes an unemployment rate that is adjusted for demographic changes. This lowers the coefficient on the minimum wage in an equation defining the NAIRU to 3.7 (standard error 1.9, p-value 5%). The coefficient remains important in economic terms but its statistical significance becomes marginal.

The issue of which specification of the price-price equation is preferable does not seem interesting. This is partly because alternative specifications support the same qualitative inference and partly because the results are part of a bigger puzzle – the difficulty in discerning *any* labor market effects in a price-price equation. For example, distributed lags of the labor share or changes in payroll taxes have negligible influence when included in the equation above. Consistent with this, price-price equations examined by other researchers (for example, Gordon, 1988) have suggested that labor costs are "irrelevant to inflation". This finding is difficult to reconcile with the apparent cointegration of prices and unit labor costs, or with the evidence presented in Appendix 1, which indicates that wages have a large effect on the price of business output, the largest component of most other price measures. Accordingly, the weakness of labor market effects is puzzling. Perhaps the equation is misspecified; perhaps there is some offsetting influence; or perhaps the effects are lost in the noise.

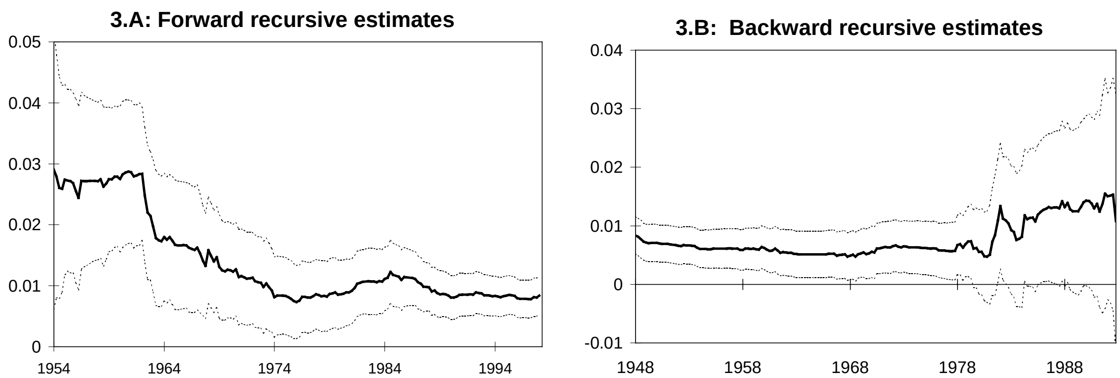
Puzzles like this make price-price equations difficult to interpret. Pending their resolution, I draw the following conclusions from this subsection. The effect of the minimum wage on the NAIRU is less clearly discernible in the price-price equation than in the wage equation. To some extent this is to be expected, given the extra layer of noise. Nevertheless, the results are consistent in qualitative terms. The evidence in this subsection corroborates rather than rejects the hypothesis that minimum wages have an important effect on the NAIRU. This hypothesis is not inconsistent with the framework most commonly used for estimating the NAIRU. I discuss price-price equations further in Appendix 1.

4 ROBUSTNESS OF WAGE EQUATION ESTIMATES

4a. Stability of minimum wage effects

Chart 3 below shows recursive estimates of the effect of the minimum wage on aggregate wages plus and minus 2 standard errors.[8] The chart indicates that the coefficient on the level of the minimum wage exceeds zero by more than two standard errors even when most observations are disregarded. Two decades of data at either the beginning or end of the sample are enough to generate a t-statistic well above 2 and a stable coefficient. At least, that is the case for estimating the effect of the minimum wage on nominal wage growth. Because the coefficient on unemployment is imprecisely estimated in small samples, the effect on the NAIRU is more sensitive to sample period.

Chart 3: Recursive coefficient on level of minimum wage
plus and minus 2 standard errors



If the significant role of the level of minimum wages were a statistical fluke then its coefficient would be unstable. The same accident is unlikely to recur in different sample periods. Furthermore, adding information should weaken the evidence of strong effects – the coefficient would decline and standard errors would fail to narrow. Chart 3 presents mixed evidence of this.

For subsamples estimated up to the early 1960s, the coefficient is significantly larger than later estimates. For other periods, the coefficient varies within a range that is

[8] Coefficients in the variance equation and on productivity and the wage-freeze are constrained to equal their full-sample estimates.

small both in economic terms and relative to the noise in the data. These breaks are reflected in Andrews-Ploberger tests, which show clear instability of the overall equation (p-values below 1 per cent) if breakpoints before the early 1960s are included, but little instability (p-values above 10 per cent) if these early breakpoints are disregarded.

A movement by a coefficient of more than two standard errors indicates that the coefficient does not measure a reliable feature of the data. This is the strongest reason of which I am aware for doubting that my estimates of the effect of the minimum wage would apply outside my sample. That said, it is unclear how heavily the early indications of instability should be weighed. The later pattern of stability may be more important.

The instability appears to be related to the limited early variation in the minimum wage. The first two within-sample increases in the minimum wage, in 1950 and 1956, were followed by large surges in wages, which are picked up by the high coefficient on the level of the minimum wage. Subsequently, as later increases in the minimum wage are followed by modest increases in aggregate wages, the coefficient declines, stabilizing around its full-sample estimate. It may be that some influence for which I have not controlled was strongly correlated with the minimum wage in early periods, but this correlation weakened over time. Candidates include the Korean War, which coincided with the first increase in the minimum wage; and a divergence between measures of wages with variable weights and measures with fixed industry weights (Gordon, 1971 p117), which coincided with the second increase. (Though Gordon (1981 p344) describes the unusual inflation of 1956 as a well-known mystery.)

Wage data from the 1940s and early 1950s are commonly regarded as unreliable. They are heavily influenced by forces that are difficult to control for, including measurement error (Gordon, 1971, pp115-117) and wars, wage controls and their aftermath (Perry, 1970, p421). For these reasons, most researchers have disregarded this period, even when degrees of freedom were scarce and the effect of the Korean War wage controls was germane. Thus unusual correlations in early data samples are not unexpected. Consistent with this, later evidence of stability suggests that unobserved influences in early subsamples may represent a temporary anomaly.

However, the difficulty in identifying these influences raises the possibility that they lie behind later more moderate estimates. Omitted variable bias is a potential problem with any non-experimental research; instability is evidence that it actually is important in this context. The most direct way of addressing this concern is to test the inclusion of a large number of omitted variables. I do this in the following section.

There is no *a priori* interest in early breakpoints. In contrast, a pertinent before-and-after comparison involves splitting the sample period at 1980. After 1980, wages are measured from the Employment Cost Index rather than from the national accounts. Also around this time the process determining the minimum wage changed: from periodic adjustment to "benign neglect". A natural concern is that these changes may affect the estimated impact of the minimum wage.

Table 2 below shows the effect the minimum wage before and after 1980. Specifically, coefficients on the level of the minimum wage and the constant are allowed to vary while all other coefficients are constrained to be the same across samples. The direct effect on wage growth is shown in column (1) and the effect on the NAIRU (allowing for an offsetting effect from unemployment) is shown in column (3).

Table 2: Effect of Minimum Wages Before and After 1980

Period	**Contribution to wage growth**		**Contribution to NAIRU**	
	Coefficient	standard error	Coefficient	standard error
	(1)	*(2)*	*(3)*	*(4)*
1948:3 to 1998:2	0.0084	0.0016	5.9	1.3
1948:3 to 1980:1	0.0109	0.0026	7.7	1.8
1980:2 to 1998:2	0.0050	0.0029	3.5	2.0

The coefficients on the minimum wage are smaller in the second sample period. However the difference is small relative to the noise in the data, with a p-value of 31%, and the coefficient remains important in economic terms. Other splits at 1980 give similar results.[9]

[9] If all coefficients (except those in the variance equation and on productivity and the wage-freeze) are

Much of the interest in the role of minimum wages comes from its substantial decline over the 1980s. The NAIRU also declined over this period (by about 1-2 percentage points according to studies cited in Section 2c). However, this 'coincidence' does not drive the results; it is what would have been expected on the basis of past relationships. As Table 2 indicates, the estimated coefficient is similar using data only up to 1980. The low frequency variations in the data are consistent with, but not necessary for, my results.

The absence of a significant change in the minimum wage coefficient before and after 1980 also implies that neither the change in wage measurement, nor the change in policy regime, fundamentally altered the effect of the minimum wage. As Hendry (1995 Ch5.9) and Ericsson and Irons (1995) argue, if the process determining one variable (in this case the minimum wage) changes[10], but the measured effect of this variable on another (average wages) does not, then that effect can be considered to be "structural" and invariant to changes in the policy process. This invariance or "superexogeneity" means that the "Lucas critique" does not apply. This property is necessary if the model is to be used for policy purposes.

4b. Sensitivity analysis

To guard against misspecification bias and as a guide to future research, Table 3 presents some tests of restrictions imposed in estimating my wage equation. To assess whether important information is potentially being ignored, the table presents a series of F-tests on additional regressors. (Usually included as a current level and one lag, to allow for both level and change effects). To assess whether this information matters, the table also shows the resulting coefficient on the level of the minimum wage and its standard error.

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allowed to change, as in Chart 3, the change in the minimum wage coefficient is slightly smaller than shown in Table 2. A Wald test for such a break has a p-value of 50%. If only the coefficient on the minimum wage is allowed to change, the coefficients are almost identical: 0.008399 versus 0.008397.

[10] As a simple indication that the process determining the minimum wage changed, I regressed the coverage-adjusted relative level of the minimum wage on its lagged value and a constant. The p-value of a break in the constant at 1980:4 is 0.2%. Such a break has a p-value of 45% in my wage equation.

For ease of interpretation, I multiply the coefficient on the level of the minimum wage and its standard error by 100. This gives an adjusted coefficient of 0.84 in the preferred specification over the full sample, with a standard error of 0.16. Over a shorter sample of say 1953:Q1 to 1996:Q4, corresponding to the data availability of some of the excluded variables, the coefficient (x100) is 0.59, with standard error 0.16.

TABLE 3: Tests of restrictions and sensitivity of minimum wage coefficient

Restriction	**P-value of restriction**	**Effect on Minimum Wage** Coefficient x100	standard error x100	**note**
I. Excluded Variables				
Aid to families	85%	0.80	0.20	a
OASDHI benefits	10%	0.76	0.17	
“other transfer payments”	8%	0.82	0.17	
total transfer payments	14%	0.80	0.16	
average tax rate	14%	0.62	0.16	
gross job destruction rate	47%	0.33	0.24	b
proportion unemployed for > 5 weeks	22%	0.77	0.16	c
mean duration of unemployment	48%	0.84	0.17	
vacancy rate	37%	0.65	0.21	d
real product wage and productivity	4%	0.76	0.17	
real consumption wage and productivity	4%	0.90	0.26	
Dummy for low wage growth	0.1%	0.91	0.16	e
import share	1%	0.93	0.16	f
variable inertia	71%	0.84	0.16	g
Price wedge	13%	0.87	0.17	h
labor force	3%	0.95	0.17	
unionization rate	80%	0.72	0.19	i
Nixon “on” and “off”	76%	0.85	0.16	j
level of payroll tax	5%	0.84	0.16	
strikes	60%	0.75	0.19	k
female labor force participation	34%	0.80	0.17	l
immigration	72%	0.62	0.16	m
import prices	35%	0.71	0.17	n
food and energy prices	19%	0.66	0.16	o
II. Specification checks				
Removal of constraints on:				
- inflation	0.2%	1.19	0.19	
- demography	51%	0.94	0.22	p
- coverage	0.01%	0.75	0.17	q
Extra lag of:				
- change in wages (constrained)	99%	0.84	0.17	r
- change in wages (unconstrained)	2%	1.14	0.21	
- unemployment rate	74%	0.84	0.17	
- change in minimum wage	89%	0.84	0.16	
- payroll tax rate	19%	0.86	0.16	

current inflation (constrained)	0.06%	0.76	0.16	s
current inflation (unconstrained)	0.01%	1.14	0.19	t
current change in unemployment	0.01%	0.88	0.16	
time and time squared	4%	0.98	0.22	
decade dummies	41%	0.89	0.20	
Unemployment above mean	1%	0.85	0.16	
III. Restrictions not imposed				
Excluding unemployment benefits	9%	0.88	0.16	
No minimum wage interaction	0.01%	0.91	0.17	
Excluding productivity	1%	0.94	0.16	
productivity coefficient = 1	2%	0.75	0.16	
IV. Non-nested respecifications	*standard error* x*100*			
Minimum wage deflated by:				
- product prices	0.241	0.61	0.11	
- consumer prices	0.241	0.61	0.11	
- average hourly earnings	0.243	0.90	0.17	
No demographic adjustment	0.244	1.01	0.17	
Stochastic productivity trend	0.247	0.81	0.20	u
No coverage adjustment	0.255	0.47	0.17	
Product prices	0.249	0.43	0.18	v
unweighted least squares	0.318	1.15	0.18	w
GARCH (1,1)	0.338	0.74	0.14	
No splicing of wages	0.396	1.19	0.23	x

Notes to Table 3:

a) This, and the next four rows represent corresponding line items from Table 2.1 of the NIPA tables, all divided by gross personal income; current and previous quarter.

b) also called the frequency of dismissal. The ratio of recent job loss of those who have not been unemployed for more than 5 weeks to total employment. Data seasonally adjusted by the Federal Reserve Board. Without this variable, the coefficient on the minimum wage (x100) is 0.43 (standard error 0.23) - the result of a substantially shorter estimation period: 1976:Q3 to 1997:Q1

c) as with mean duration, this was lagged twice, matching the lag on unemployment. The proportion unemployed over 15 and 27 weeks gave even weaker results

d) help wanted index divided by civilian employment; lagged once and twice

e) a dummy equal to 1 in the 23 (out of 200) quarters in which fitted wage growth is less than 0.75 per cent a quarter. The dummy boosts wage growth by 0.14 percentage points.

f) ratio of imports of goods and services to gross domestic purchases. Current and previous quarter

g) allowing the coefficient on last year's inflation to be a linear function of the inflation rate.

h) Quarterly wage growth is *lower* by 0.16 percentage points for every percentage point that the lagged 4-quarter change in product prices exceeds that in consumer prices. Inclusion of this variable is equivalent to adding product prices while preserving the inflation neutrality restriction.

i) union members as a share of civilian employees. Rate for current and previous year; 1950 - 1995

j) Gordon's incomes policy dummies. The coefficient and standard error on the 1971 wage freeze is not affected by inclusion of these variables.

k) Percentage of time lost due to stoppages; current and previous year

l) female labor force divided by non-institutional civilian population, from MPS data base; current and previous quarter

m) immigrants admitted divided by U.S. population, from Statistical Abstract of the United States. Current and previous year for 1951-1994.

n) Chain-weighted price index for imports of goods and services divided by that for non-farm business output. Current and previous quarter

o) Difference between the chain-weighted price indexes for personal consumption expenditures including and excluding food and energy.

p) allowing the unemployment rate and the demographic adjustment to enter separately

q) allowing the minimum wage and coverage to enter separately

r) lagged dependent variable with coefficient plus those on inflation terms summing to one (preserving inflation neutrality)

s) contemporaneous first difference in consumption prices, with coefficients on lagged inflation summing to one
t) as above, without constraint
u) measuring trend productivity growth as a 10 year back average. This gives a similar coefficient and fit in the wage equation as the deterministic trend, but does not perform as well in the price equation.
v) using deflator for non-farm business output instead of for consumption expenditures, for 1952:2 to 1998:2. For this period the baseline specification has an equation standard error of 0.2261
w) restricting the coefficient on time in the variance equation to zero (actually, this is nested, with a p-value < 0.01%)
x) using compensation-per-hour (not spliced with the ECI) for both the dependent variable and the denominator in the minimum wage term. Estimation by unweighted least squares.

As a general observation, relaxation of exclusion or other restrictions makes less difference than changes in sample period. Most changes in specification alter the coefficient on the minimum wage by less than a standard error. Large reductions in the coefficient occur if the minimum wage is not adjusted for coverage, if product prices are used instead of consumer prices or if data availability substantially shortens the estimation period. Variations such as these can be interpreted as disregarding relevant information – this can alter the results, but doing so is uninteresting. Even so, the coefficient on the minimum wage is always positive and important in economic terms. The inference I draw from the results above (and many more that are not reported) is that the quantitative effect of the minimum wage on nominal wage growth is sensitive, but not unduly so, to variations in model specification. The qualitative effect is robust.

The estimated effect of minimum wages on the NAIRU is less robust than its effect on wage growth, reflecting sensitivity of the coefficient on unemployment to alternative specifications. However, sensitivity of the unemployment effect does not seem an interesting issue – given that my estimates are in line with out-of-sample evidence from other countries and time periods. The chief uncertainty concerns the question of whether minimum wages are an important determinant of wage growth, rather than the subsequent issue of the amount of unemployment required to offset this. The focus of my sensitivity analysis reflects this.

Although many restrictions in the table are interesting, two sets of these warrant a special comment. I discuss some other restrictions in Tulip (2000).

It might be thought that the minimum wage simply serves as a proxy for deeper social forces, such as readiness to intervene in markets or egalitarian attitudes. However, if this were so, then other elements of the social safety net should be important. However, as indicated in the first five rows, they are not; nor do they affect the significance of the minimum wage.

The existence of a unique NAIRU is a simplifying assumption rather than an implication of my data set. The restriction that the coefficients on lagged inflation sum to one is significantly rejected, as are other exclusions that would relax this constraint (such as inclusion of squared fitted values or the unconstrained lagged dependent variable). Coefficients summing to significantly less than one could arise because of the existence of a long-run tradeoff between inflation and unemployment, because inflation is expected to revert to its mean or because of errors in the measurement of prices. My results could easily be interpreted in terms of the effect of the minimum wages on a (steep) long-run tradeoff. But the implications would be similar, and the story would be more complicated and less interesting to most readers.

5 CONCLUSION

A significant effect of the minimum wage on the NAIRU is evident in four distinct data sets: wage growth from 1948 to the 1970s; wage growth from the 1970s to 1998; the behavior of prices; and international comparisons of changes in the NAIRU. These different sources of information are consistent with a 10 per cent increase in the relative minimum wage raising the NAIRU by about half a percentage point. Such an effect is also consistent with – and explains – time-varying estimates of the NAIRU.

Doubt about the implications of my wage equation arises from an apparent structural break in the early 1960s – when the coefficient on the minimum wage dropped by over two standard errors. This could be disregarded because early wage data are poor. Or it could be taken as evidence of some unobserved influence that was correlated with the minimum wage. Both of these possibilities could explain away my other results. However, their effect would need to be similar in each of the data sets noted above, which may seem unlikely. Furthermore, my efforts to find omitted effects on wages were unsuccessful. The other sources of information can also be challenged. For example, the coefficients in my price-price equation are imprecise and my cross-country comparisons may not survive multivariate analysis.

All empirical research is subject to doubt. Nevertheless, a wide range of evidence indicates that the minimum wage has an important effect on the NAIRU.

Appendix 1: Price equation estimates

The price equation plays a secondary role, mainly serving to establish that prices mimic wages, and hence that the NAIRU can be estimated by the wage equation. Estimates of an equation similar to (2) – without the simplifying restriction that the coefficient on the change in labor costs is unity – are presented below. The dependent variable is the change in the logarithm of the chain-weighted price index for the business sector excluding agriculture and housing. This is the broadest measure of prices for which wages and productivity can be measured on a consistent and reliable basis. Accordingly, it is an appropriate measure for assessing how prices respond to changes in unit labor costs. It also has less measurement error than other price series, permitting more confident inferences. The business sector, excluding housing and agriculture, constitutes about three-quarters of GDP. So its prices will have a dominant influence on other economy-wide price measures.

Table 4: A Price Markup Equation

Regressor	Coefficient	Standard error
contemporaneous change in unit labor costs (instrumented)	.22	.061
average change over previous four quarters in unit labor costs	.55	.066
ln(price / unit labor costs) (-1)	-.119	.022
unemployment rate(-1)	-.00049	.00014
average change over last 8 quarters in relative energy prices	.034	.0080
change in relative import prices (-2)	.050	.013
change in relative farm prices	-.019	.0026
dummy for Nixon price controls	-.015	.0033
Constant	.53	.098

sample	1955:Q1 1998:Q1
standard error	0.0022
R-Squared	.91
RESET test (with squared fitted values)	4%
Breusch-Godfrey test for 4^{th} order serial correlation	7%
White's test for heteroskedasticity/misspecification	23%
Jarque-Bera test for normality	52%
Andrews-Ploberger test for stability (1957:Q2 - 1996:Q1)	7.6 (p >10%)

All change terms represent log differences. The wage and unemployment series are the same as used in the wage equation. The trend productivity term is the same, but without the moving back-average. Unit labor costs are wages divided by trend productivity. I instrument the contemporaneous change in unit labor costs using predicted values from the wage equation (though OLS estimates are similar). Relative price terms subtract the average change in the dependent variable over the previous eight quarters.

The equation suggests that prices are a markup on unit labor costs, with changes in unemployment and in energy, import and farm prices causing temporary fluctuations about a stable long-term relationship. The error-correction term is highly significant, in both statistical and economic terms. About 12 per cent of a deviation of prices from their long run level is eroded each quarter - implying that half the deviation disappears within six quarters.

Because of its secondary role, I do not subject the price equation to the battery of tests I perform on the wage equation. However, one exclusion restriction is worth mentioning. When the level of the minimum wage is included, it has a negative coefficient with a p-value of 15%. This effect is not robust (for example, the coefficient changes sign if the sample is shortened slightly and a time trend is included). Nevertheless, it suggests that the minimum wage depresses price margins, as implied by monopsony models. This will offset the inflationary impact of the minimum wage in the short run. It may explain why the effect of the minimum wage is not easier to discern in a price-price Phillips curve (section 3b). However, the error-correction coefficient is not affected. Profit margins do not fall without limit. As they approach their new lower level, prices resume tracking nominal unit labor costs (which, in turn, depend on the minimum wage). Monopsony means a one-off change in the price level, not an ongoing change in inflation. The long run effect of the level of the minimum wage on the NAIRU is not affected.

Viewing prices as a markup differs from the popular approach of explaining inflation through a price-price Phillips curve, where labor costs are replaced by lags of the dependent variable. This approach has been partly inspired by Gordon (1988) who estimated equations in which productivity is measured for the business sector, but prices relate to GNP. These series are weakly correlated across sectors which may explain why unit labor costs appear to be "irrelevant for inflation".

When these variables are measured for the same sector, as above, then unit labor costs outperform lagged inflation in explaining current inflation. The addition of four lags of inflation to the markup equation is jointly insignificant, with a p-value of 37% and coefficients summing to 0.16 (standard error 0.09). In contrast, the unit labor cost terms remain important, with the coefficients on the change in unit labor costs summing to 0.61 (standard error 0.10) and that on the level of unit labor costs rising to 0.138 (standard error 0.025). An alternative non-nested specification including 24 lags of the dependent variable, with coefficients constrained to lie on a 3^{rd} degree polynomial (the unconstrained sum of these coefficients is 1.03), the same controls for energy, import and farm prices, but excluding the wage terms, has an equation standard error of .0025, 15% larger than the equation set out above. Inclusion of the level and change in unit labor costs remain highly significant, with p-values less than 0.01 per cent.

The irrelevance of the lagged dependent variable when unit labor costs are appropriately controlled for implies that lagged inflation does not play a direct causal role in determining prices. Although they do not reflect behavioral relationships, price-price equations are nevertheless useful for forecasting price measures for which productivity cannot be measured on a comparable basis. This is because unit labor costs for a particular sector can often be proxied more accurately by lagged inflation than by unit labor costs for the business sector.

Appendix2: Data

The data for my three main regressions can be downloaded as a Eviews 3.1 file from http://petertulip.homepage.com

Table 5: Data Measures and Sources

Variable	Measure	Data source
Wages from 1980:Q1	Compensation of non-farm private industry workers	Employment Cost Index, BLS
Wages to 1980:Q1	Compensation per hour; non-farm business sector	unpublished BLS data, from "Productivity and Costs"
consumer price from 1947:Q1	Price index for personal consumption expenditure	NIPA Table 7.1
consumer price 1942 to 1947	Implicit price deflator for personal consumption expenditures at 1987 prices; interpolated from annual data	BEA (1993); Table 7.1 row 16
product price	Price index for output of the non-farm business sector, excluding housing	NIPA Table 7.14
Productivity	Output per hour of all persons in non-farm business sector	"Productivity and Costs" BLS
unemployment rate	unemployment rate with constant demographic weights, adjusted for breaks in series.	Federal Reserve Board of Governors
demography	Difference between above and civilian unemployment rate, adjusted for breaks in series.	Federal Reserve Board of Governors
unemployment benefits	Unemployment insurance benefits divided by unemployed civilians aged 16 and over, divided by wages	NIPA Table 2.1 (row 17); Employment and Earnings, Table A-1
minimum wage	statutory minimum by state, divided by wages	see section 3 of this appendix
Minimum wage coverage	proportion of non-supervisory employees covered by FLSA	Ehrenberg and Smith (1996 p118)
payroll taxes	employer's statutory OASDHI contribution rate (seasonally adjusted from 1980:Q1).	Social security Administration (1998)
wage freeze	1 in 1971:Q4; -0.6 in 1972:Q1	FRB/US database (DWPC)
food and energy prices	difference between price indexes for personal consumption expenditures with and without food and energy	Federal Reserve Board of Governors
import prices	deflator for imports of goods and services excluding oil, computers and semiconductors	Federal Reserve Board of Governors
farm prices	price index for agricultural output	FRB/US database (PH.GDP_D87YAGF)
energy prices	price of oil to 1956 spliced with composite of prices for oil, coal and gas.	FRB/US database (PCENG, POIL)
price controls	1 from 1971:Q3 to 1974:Q1; -3.67 from 1974:Q2 to 1974:Q4	Federal Reserve Board of Governors

Data construction

Wages, prices, productivity, unemployment benefits and the minimum wage are measured in logarithms.

For data published in both seasonally adjusted and unadjusted form, the adjusted series is used.

All changes (including of the payroll tax rate and the unemployment rate) are first differences.

"Last year's inflation" is [$\ln price_{t-1}$ - $\ln price_{t-5}$] / 4

"Previous four years" inflation' is [$\ln price_{t-5}$ - $\ln price_{t-21}$] / 16

"Trend productivity" is a linear trend fitted to the logarithm of output per hour of all persons in non-farm business sector, with a kink at 1973:Q1. For the measure used in the wage equation, I then take a 5-year back average of the quarterly change in trend productivity.

All price series, except those described as deflators, are chain-weighted.

The price wedge is the average difference between the log quarterly change in consumer prices and product prices. This is zero to 1981 then 0.0011 thereafter.

The demographically adjusted unemployment rate is the average of unemployment rates for five age-sex categories, weighted by shares of 1993 unemployment, as a percentage. The series is adjusted for several breaks, the largest of which are a permanent increase of 0.08 percentage points in the unemployment rate in January 1994 due to new survey design and a gradual increase from 1980 to 1990 of 0.1 percentage points arising from undercounting that was rectified with the 1990 Census benchmarks.

Measurement of the minimum wage

Although I refer to "the" minimum wage, it is not uniform. Two variations for which I make allowance are the different rates specified by some states and the incomplete coverage of the Federal legislation.

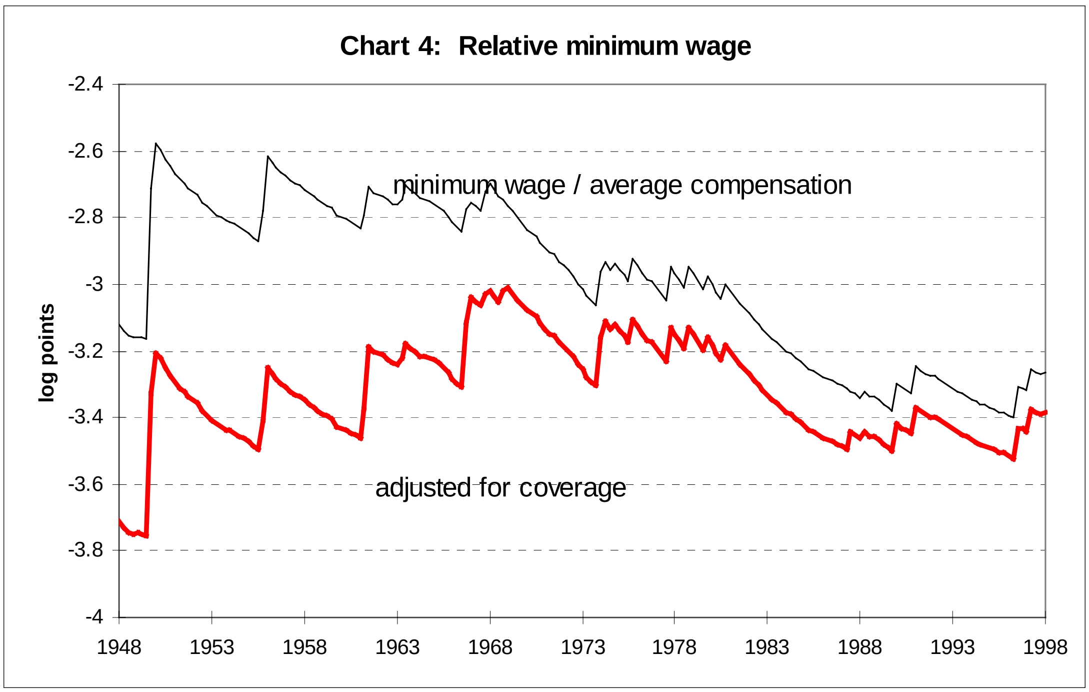
I use whichever is the higher of the main state and Federal rate for each state for each month, and weight each state by 1996 employment shares. A time series for the Federal minimum wage, and much other relevant information, is available at the Employment Standards Administration web site: http://www.dol.gov/DoL/ESA/public/ minwage/ main.htm. Data on state minima for 1950 to 1981 come from Questor (1981, Table 1a). Data from 1981 to 1996 come from the Council of State Governments (1998). Current state rates are available at the ESA web site. When legislation specifies different minimum wages for different workers, I use the same rate as Neumark and Wascher (1992 p59) where available, otherwise the rate that applies to adult males in most industries (typically the Federal rate). Prior to 1950, I assume that no state minimum wages are binding (in 1950 only 6 states had minimum wage laws that applied to men).

The state-adjusted series exceeds the Federal minimum by 4 per cent in 1974 and 6 per cent in 1990, but otherwise is barely distinguishable from it. Between 1950 and 1998 the divergence between the series averages 0.7 per cent.

The importance of minimum wage legislation depends on how much of the economy it covers. To allow for this, I multiply the minimum by the proportion of non-supervisory employees covered by the Federal FLSA minimum wage provisions. Federal coverage seems relevant given the dominant role of the Federal rate. Ehrenberg and Smith (1996 p118) present estimates (from the Employment Standards Administration) of coverage at the time of each legislative change since 1938. Although this series has problems, such as an undated increase in the 1980s unrelated to legislation, a similar independently compiled series in the FRB/US database gave almost identical results. I assume, for convenience, that coverage remains constant between legislative changes. It would be more accurate to use coverage data for each year, though this would not greatly change the results.

I expect that further data adjustments would make small changes to the quantitative results, without affecting the main conclusions. In this regard, the largest adjustment would probably involve the phased introduction of the minimum wage to newly covered workers in 1961, 1966 and 1974. Whether this or other adjustments are appropriate is unclear, given that many exemptions, such as those for youth, appear not to be utilized (Card and Krueger, 1995).

Chart 4 below shows the logarithm of the minimum wage, divided by average hourly compensation, and this series adjusted for coverage.



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