Chapter 1:

Do minimum wages raise the NAIRU ?

Summary

Probably yes.

The level of the minimum wage (relative to average wages) has a significant effect on nominal wage growth and hence on inflation. This can be offset by extra unemployment; so the minimum wage increases the Non-Accelerating Inflation Rate of Unemployment or NAIRU.

This effect is 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.

1.1 Introduction

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

In this chapter, 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. On average, a 10 per cent increase in the minimum wage, relative to average wages, seems to raise the NAIRU by 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. The surprisingly favorable behavior of inflation and unemployment in the USA over the 1990s can be attributed, in part, to the low level of the minimum wage.

Although the effect of the minimum wage on the NAIRU is important in macroeconomic terms, it is perhaps more directly relevant to labor market policy. Specifically, it measures the sustainable unemployment arising from the minimum wage. This idea has been neglected in previous research. Surveys of the consequences of the minimum wage by Brown, Gilroy and Kohen (1982) and Card and Krueger (1995, Chapters 6 and 7) focus on changes in labor demand. However, this effect might be offset by Phillips curve effects; that is, the tendency of unemployment to put downwards pressure on wages. If unemployment arising from the minimum wage represents a movement along an unchanged Phillips curve (that is, if this unemployment exerts as much downward pressure on wages as unemployment arising from other reasons), it would imply continually falling wages. This would increase employment elsewhere and the unemployment would be temporary. Alternatively, if the minimum wage shifts the Phillips curve (as I estimate it to do) there might be a sustained increase in unemployment that exceeds any change in labor demand.

The focus on labor demand effects could be justified if interest lay in non-employment rather than unemployment and if there were no Phillips curve effects. But policy makers seem to be much more interested in unemployment, and the Phillips curve is a well-documented empirical regularity, even if its precise form is controversial.

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.

A larger body of research uncovers similar empirical relationships, without explicitly interpreting these as an effect of the minimum wage on the NAIRU. For example, Adams (1989) notes that the gap between average and minimum wages is significant in a wage equation. Card and Krueger (1995 p163) suggest that increases in the minimum wage may gradually flow throughout the wage structure. And, as I discuss in Section 1.3.1, many researchers have noticed an apparent “tradeoff” between inequality and unemployment in cross-country comparisons. My work adds to these studies by showing that the relationship between the minimum wage and aggregate nominal wage growth is a robust empirical regularity that explains variations in the NAIRU.

The plan of this chapter is as follows. Section 1.2 defines the NAIRU, explains how estimates of it can be derived from a wage equation and presents those estimates. Section 1.3 shows that the effect of the minimum wage on the NAIRU is also evident in the behavior of prices and in international comparisons. Section 1.4 looks at the estimates from the wage equation in detail and shows that they are robust to variations in model specification and sample period. Section 1.5 concludes.

As the plan indicates, the focus of this chapter is empirical. It seems appropriate to determine whether minimum wages have an important effect on the NAIRU before asking how or why this might be so. I discuss how this effect might be interpreted in Chapter 3 and develop a formal model in Chapter 4. But because there is some interdependence between these questions, it may be useful to briefly note some possible channels of influence here.

The minimum wage will increase the NAIRU if it reduces the demand for unskilled workers or increases workforce participation, *and* if this extra unemployment puts little downward pressure on wages. Higher unemployment would be associated with a similar wage outcome.

The minimum wage will also increase the NAIRU if it causes other nominal wages to gradually increase, for a given rate of unemployment, and if these wage increases flow into higher prices. Common reasons for this preservation of relativities include fairness and a desire to recruit, retain and motivate workers. Pressure on other wages could also occur if the minimum wage acts as a safety net, allaying worker insecurity. When the minimum wage is low, other wages have further to fall if a worker loses his job. This danger may cause workers to bargain less aggressively.

In Chapter 3 I review evidence on these potential interpretations. I conclude that effects associated with direct maintenance of wage relativities and labor demand might be present, but they do not appear to account for the overall effect. The evidence of effects operating through worker insecurity is stronger. For example there are many indications that increased worker insecurity accounts for a large part of the decline in the NAIRU over the last two decades, and this, in turn, can be attributed to the low minimum wage. The evidence on worker insecurity corroborates the macroeconomic evidence on the effect of the minimum wage that is the focus of this chapter.

1.2 Framework and Estimates

1.2.1 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 1.3.2, 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 seem to omit important influences, 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 wages 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. Earlier versions have been presented in Ando, Brayton and Kennickell (1990) and 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 reflect the main empirical determinants of 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. My attempts to do so are discussed in Chapter 4.

The wage equation is important because, in the medium to longer run, prices are assumed to mimic wage movements. This assumption seems to be realistic. As I discuss in Appendix 1 of this chapter, 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 residual. To simplify the exposition I ignore lags and constrain the coefficient on the change in unit labor costs to 1. 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 minimize 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, $P_p \approx P_p^*$. 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 arguably could be included in the specification. Using these conditions and differencing (3) gives a long-run version of equation (2):

$$\Delta P_p = \Delta W - \Delta\, prod \qquad (4)$$

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

$$\Delta P_c = \Delta P_p + wedge \qquad (5)$$

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, differential productivity movements 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. In Section 2.1, I describe the wedge in more detail.

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 a simpler, more easily interpretable equation. 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:

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

According to this definition, 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 the price *wedge* as exogenous. This approach means that 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.

1.2.2 Wage equation estimates

An econometric equation explaining nominal wage growth is set out below. 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 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 seems to reduce measurement error, permitting more precise estimates and more robust inferences than would be possible from the first measure alone. 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) and Section 2.5. My estimates are as follows (with standard errors in square brackets under the respective coefficient):

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

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

$$
\begin{aligned}
\log(\sigma^2) = \quad & -10.36 \quad - \quad .0155 \text{ TIME.} \\
& \quad \textit{[0.19]} \qquad \textit{[.0017]}
\end{aligned}
$$

sample	1948:3 1998:2
standard error	0.002435
R-Squared	.88
RESET test of functional form	1%
Breusch-Godfrey test for 4^{th} order serial correlation	17%
White's test for heteroskedasticity/misspecification	9%
Jarque-Bera test for normality	56%
Andrews-Ploberger test for stability (1951:3 – 1995:3)	13.7 (p <1%)
Andrews-Ploberger test for stability (1963:3 – 1995:3)	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 the average over the last five years of the quarterly change in trend log business output per hour, with a kink at 1973:1. "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%. Summary statistics and diagnostics are based on weighted residuals (that is, divided by the estimated standard error). Most statistics are calculated directly by Eviews[1].

[1] Version 3.1 With the following exceptions. RESET tests the addition of the squared fitted values to the equation. Breusch-Godfrey tests 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

Diagnostics are reported as p-values of F-statistics: that is, the probability of observing the sample under the null hypothesis of no problems, except for the Andrews-Ploberger test of parameter stability, which has a non-standard distribution. I discuss tests of instability in Sections 1.4.1 and 2.6 and mis-specification in Section 1.4.2.

Compared to other wage equations, my equation most closely resembles that of the MPS model, as published in Ando and Brayton (1995). It differs in many ways; the most important of which are the treatment of minimum wages, productivity and heteroskedasticity. The result of these changes is that my equation gives a better fit to a longer range of data with fewer estimated parameters. The standard error of the equation is 0.0024 log points (when weighted, or 0.0033 unweighted), compared with 0.0037 for Ando and Brayton. Most of this improvement comes from redefining the dependent variable (which seems to reduce measurement error) and inclusion of the minimum wage. These factors are partially offset by increasing the sample period (so temporary correlations are less likely to survive) and inclusion of noisy data from the 1940s and 50s.

1.2.3 NAIRU estimates

The definition of the NAIRU in equation (7) could be interpreted as including temporary influences such as the changes in payroll taxes, the residual or the different lags in the 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

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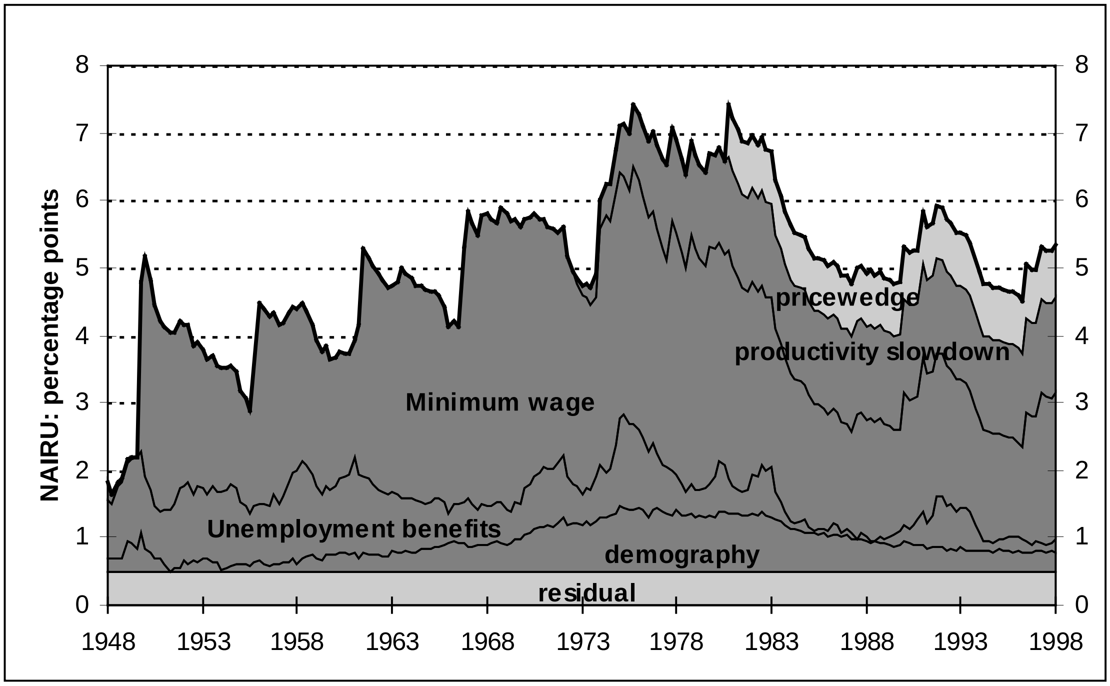
statistic both for all feasible breakpoints and for a large range over which it is insignificant. A high Andrews-Ploberger statistic implies instability. Critical values, as tabulated by Andrews and Ploberger for short periods of time. For example, to offset a percentage point increase in the social security contribution rate requires an extra percentage point of unemployment be maintained for 6 (= .88/.0014 x 1%) quarters. However, because such influences do not last, or are soon reversed, they have a small total effect. If included in the NAIRU, they would raise it by an average of half a percentage point over the full sample. (They are measured as deviations from zero, rather than from their mean).

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. Accordingly, Figure 1.1 below shows estimated contributions to the NAIRU from the constant, the level of the minimum wage, unemployment benefits, demography, productivity and the price wedge since 1948. I measure the wedge by estimating a linear trend in the log of the ratio of consumer prices to product prices, with a kink at 1981, then differencing. As discussed in Section 2.1, this approach removes transient influences such as the effect of weather on farm prices. 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 and the price wedge to be unimportant.

(1994) for their parameters of $p = 9$ and $\lambda = 7$, are 11.4 (1%) and 8.1 (10%).

Figure 1.1: Persistent Contributions to US NAIRU



To graphically stack the separate contribution of each variable I take deviations from the minimum value observed over the period. This means that the constant term in the figure, labeled “residual”, may be loosely interpreted as the NAIRU that would obtain if the other factors were at their lowest level (a kind of best-case scenario). Of course, that involves extrapolating estimates to extreme points in the sample, so this estimate is sensitive to functional form assumptions.

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

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

$$+\ \text{wedge}/\underset{[0.002]}{0.0014} + \text{demography} \tag{8}$$

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 figure). Both variables are in logarithms; so, for example, a 10 per cent increase in the minimum wage increases the NAIRU by 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. In Section 2.4, I estimate that a 95% confidence interval for the NAIRU at the end of my sample period, 1998:2, extends from 4.8 to 5.6 per cent. This is even narrower than the standard error of the constant because of negative covariances between coefficients. Further measures of uncertainty could be calculated to take account of unmodeled heteroskedasticity (Davidson and MacKinnon, 1993, Ch16.3) or serial correlation (Andrews, 1991). Although of academic interest, these sources of uncertainty seem to be unimportant relative to uncertainty about whether the model is correctly specified. I discuss specification uncertainty in Section 1.4.2.

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 directly, including Tobin (1980 p58), Stiglitz (1997, p6), Gordon (1997), Staiger, Stock and Watson (1996), the FRB/US model (1999) and the OECD (1996b, Table 1). There are many differences between (and within) these studies however a common finding is 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.[2] These results corroborate the main patterns in Figure 1.1. Perhaps more interestingly, my decomposition provides an explanation of the other researchers' results.

1.2.4 The magnitude of minimum wage effects

Were an increase in the relative minimum wage of 0.1 log points (approximately 10 per cent) 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; indeed, it is less than the rounding error in the data. However, it is substantial relative to the effect of unemployment on wages, which is the metric that is relevant to policy. 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. President Clinton has proposed increasing the Federal minimum wage from its current level of $5.15 an hour to $5.65 on September 1, 1999 and $6.15 on September 1, 2000. I estimate that this would raise the NAIRU by one percentage point, relative to a policy of no change in the nominal minimum.

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[2] 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.

As Figure 1.1 indicates, the reduction in the relative level of the minimum over the 1980s accounts for a reduction in the NAIRU of about 2 percentage points. This effect would have been greater but that coverage increased and several States increased their minimum wages above the Federal rate. Increases in the minimum in 1996 and 1997 raised the NAIRU by half a percentage point. The reduction in the minimum wage from its peak in 1969 to its latest trough, in 1996, reduced the NAIRU by 3 percentage points.

Freeman (1996 p645) argues that the minimum wage cannot be important for aggregate wage determination because few workers are directly bound by it. In the limited sense that immediate effects on wages are small, this might be correct. But then the same would be true of unemployment's effect on wages, an effect that is central to many policy discussions. And it would be true for the same reason: the proportion of the workforce that is unemployed is similar to the proportion of the workforce on the minimum wage.

Although it may seem small, the effect of the level of the minimum wage is 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 (the work of Kremers, Ericsson and Dolado (1992) suggests this assumption is reasonable), 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. In Section 2.4 I estimate a 95% confidence interval for this coefficient spans 3.6 to 9.0.

Given the novelty and importance of these results, I examine additional information on the effect of the minimum wage on the NAIRU in Section 1.3 and the sensitivity and stability of my wage equation in Section 1.4.

1.3 Corroborating Evidence

1.3.1 International comparisons

In many other countries, governments have strong influence over 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 measuring and comparing these interventions is through the ratio of the 10^{th} percentile of the wage distribution to the median. In the U.S., 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 10^{th} and 50^{th} 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. 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).

The effects of government interventions can be gauged by time-varying estimates of the NAIRU prepared 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).

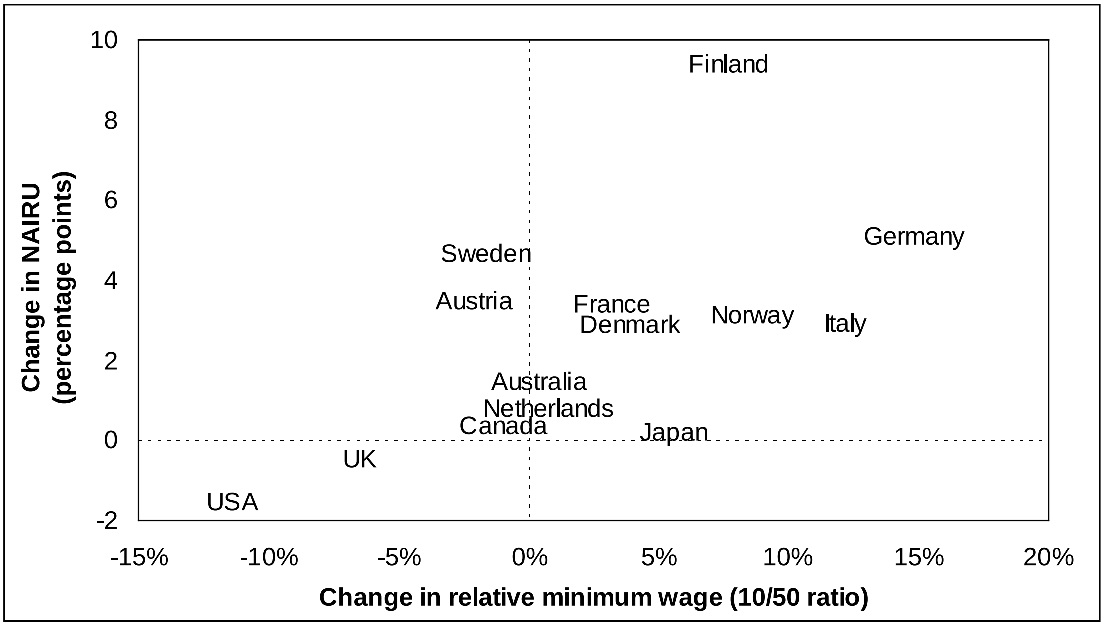
To be convincing, the effect of minimum wages on the NAIRU should be embedded in a multivariate framework. That is an advantage of a close focus on one country's data. However, models with which this could be done for other countries have not yet been developed. Nor has the literature on the determinants of the NAIRU indicated any effects 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. Figure 1.2 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, such as 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.

Figure 1.2: Changes in relative wages and the NAIRU

OECD economies; 1980 to 1995[3]



The international comparisons convey the same message as the U.S. time series: when wages at the bottom of the distribution are compressed, the NAIRU usually rises. The relationship is also similar in quantitative terms. Figure 1.1 indicated that if wages at the bottom of the distribution in the United States had kept pace with wages at the middle over this period then the NAIRU in 1995 would have been about 2 percentage points higher than in 1980. The experience of other countries, as measured in the figure 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

[3] Data relate to persons, 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 the nearest available estimate.

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 on unemployment in order to explain away this correlation.

Variations on the relationship shown in Figure 1.2 have been noticed by many observers. The OECD Jobs Study (1994, Figure 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 much simpler explanation of the relationship between unemployment and inequality is suggested by my examination of the U.S. time series. That is, the relationship may be causal. Part of the increase in unemployment in Europe may be due to attempts to raise the relative wages of low-paid workers.

1.3.2 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, my interest lies more 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 1-1: A Price-Price Equation

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	1952:3 1998:2
standard error	0.0020
R-Squared	.91
RESET test (with squared fitted values)	66%
Breusch-Godfrey test for 4^{th} order serial correlation	3%
White's test for heteroskedasticity/misspecification	99%
Jarque-Bera test for normality	61%
Andrews-Ploberger test for stability (1954:2 – 1996:4)	11.5 ($p < 1\%$)
Andrews-Ploberger test for stability (1958:2 – 1996:4)	4.6 ($p > 10\%$)

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.

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 extra lag from costs to prices. 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). The estimate from the price equation is less precise than that from the wage equation. Commensurate with this, it is more sensitive to variations in sample period and specification. Inclusion of other variables can reduce both the magnitude and statistical significance of the coefficient.

The imprecision of the coefficient on the minimum wage is part of a bigger puzzle. Given the extra layer of noise, the effect of labor market variables on consumer prices should be harder to detect than their effect on wages. Even so, these effects are surprisingly difficult to see. For example, distributed lags of the labor share or changes in payroll taxes have negligible influence when included in the equation above. More germane is the weakness of contributions to the NAIRU. To see this, denote the NAIRU presented in equation (7) and Figure 1.1 as $NAIRU^w$ and include it in place of the minimum wage, then estimate:

$$\Delta P_c = \Delta P_{c(-1)} + \mathrm{a}_1(Unemployment - constant - \mathrm{a}_2 NAIRU^w) + other \qquad (9)$$

The estimated coefficient on $NAIRU^w$, a_2, is 0.58 (standard error 0.20), significantly less than 1. This implies that contributions to the NAIRU estimated within the wage equation have slightly more than half as large an effect when included in the price equation. This equation fits the data slightly worse than the specification in Table 1-1, suggesting that elements of the $NAIRU^w$ other than the minimum wage may be erroneously included. Measurement error bias presumably accounts for some of the quantitative discrepancy.

More dramatically, some researchers, most notably Gordon (1988), have failed to find a significant effect of any labor costs in equations like that presented above. They have concluded from this that wage behavior is irrelevant to inflation. This view is difficult to reconcile with the evidence presented in Appendix 1, that wages have a large effect on product prices, which dominate most price measures. Yet the apparent 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 discernable in the price-price equation above than in the wage equation. However, the results are consistent in qualitative terms. The evidence in this sub-section corroborates the hypothesis that minimum wages have an important effect on the NAIRU. More convincingly, it indicates that this hypothesis is not rejected within the framework most commonly used for estimating the NAIRU. I discuss price-price equations further in Appendix 1.

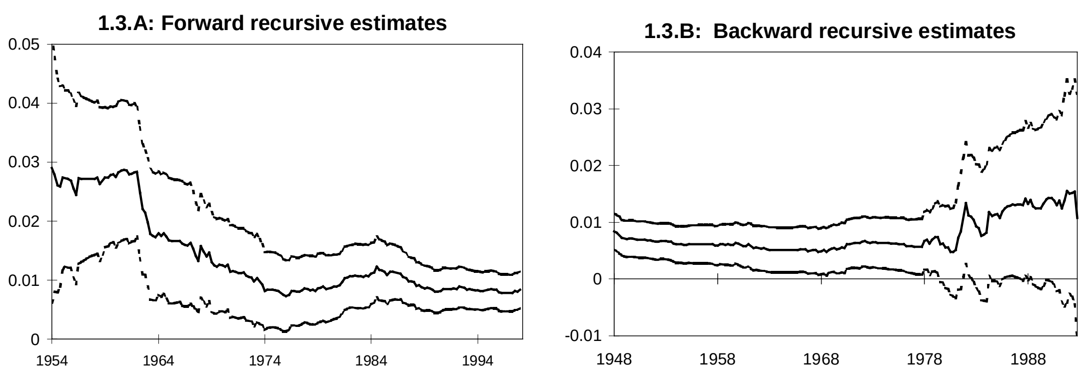
1.4 Robustness of Wage Equation Estimates

1.4.1 Stability of minimum wage effects

In Section 2.6 I comment on the overall stability of the wage equation. In this section I focus on the coefficient on the level of the minimum wage.

Figure 1.3 below shows recursive estimates of the effect of the minimum wage on aggregate wages plus and minus 2 standard errors. (For these, and other sub-sample estimates in this section, I constrain the coefficients in the variance equation and on productivity and the wage-freeze to equal their full-sample estimates). The figure 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.

Figure 1.3: Recursive coefficient on level of minimum wage plus and minus 2 standard errors



If the significant role of the level of minimum wages was a statistical fluke then its coefficient would be unstable. The same accident is unlikely to recur in different sample periods. Furthermore, adding extra information should weaken the evidence of strong effects – the coefficient would decline and standard errors would fail to narrow. Figure 1.3 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 small both in economic terms and relative to the noise in the data. This is reflected in Andrews-Ploberger tests, which show clear instability of the overall equation (p-values well 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.

Andrews-Ploberger tests may be biased towards finding instability when residuals are heteroskedastic and regressors are stochastic.[4] However, given the wide margin by which the test rejects, this bias would need to be large for the apparent instability to be confidently attributed to sampling variability. Residual autocorrelation is also unlikely to be the cause, given that p-values from Breusch-Godfrey tests in early subsamples typically exceed 20%. Nor did experimentation with alternative functional forms and specifications account for the instability.

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

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[4] Greene (1995 p355) notes several studies of the tendency of Wald breakpoint tests to over-reject. Hendry (1995, Figure 6.8d) provides Monte Carlo evidence that stability tests of models with dynamic regressors over-reject.

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 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 albeit to a lesser extent. 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.

A common response to potential instability is to ignore the early data. This may be appropriate if simplicity is at a premium and the focus is on short-term relationships (as with forecasting). However, a practical difficulty with this is the absence of a clear dividing line between "bad" early data and "good" late data. My attempts at modeling the heteroskedasticity in the residuals suggest that the data quality improves gradually over time. More importantly, policy analysis that disregards uncomfortable evidence is unconvincing. Extra information is particularly valuable for assessing correlations among persistent variables, of which independent observations are few. The likelihood of finding large spurious correlations is reduced as sample size increases. (Although t-statistics increase, bias declines.) Accordingly, I use an unusually long data set, extending back to 1948, but give earlier observations a diminished weight, with the weight being determined by the noise in the data.

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 1-2 presents estimates of the effect of minimum wages before and after 1980. 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). Comparing rows (2) and (3) shows the effect of re-estimating the entire equation for each subsample. Comparing rows (4) and (5) shows the effect of letting the coefficients on the minimum wage and constant vary, while constraining all other coefficients to be the same across samples. (If only the coefficient on the minimum wage is allowed to change, the coefficients are identical: 0.008399 versus 0.008397. This apparent stability is arguably an artifact of the procedure).

Table 1-2: Effect of minimum wages for different periods

Period	**Contribution to wage growth**		**Contribution to NAIRU**	
	Coefficient	**standard error**	**Coefficient**	**standard error**
	(1)	*(2)*	*(3)*	*(4)*
(1) 1948:3 to 1998:2	0.0084	0.0016	5.9	1.3
allowing all coefficients to vary (p-value of Wald breakpoint test: 50%)				
(2) 1948:3 to 1980:1	0.0090	0.0028	9.8	3.7
(3) 1980:2 to 1998:2	0.0062	0.0034	3.3	1.7
only coefficients on minimum wage and constant vary (p-value of F-test: 31%)				
(4) 1948:3 to 1980:1	0.0109	0.0026	7.7	1.8
(5) 1980:2 to 1998:2	0.0050	0.0029	3.5	2.0

In both cases 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 and the qualitative implications of the model do not change. This has several implications.

The absence of a significant change in the coefficient on the minimum in 1980 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[5], 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 these policy interventions. 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.

Much of the interest in the role of minimum wages comes from its substantial decline over the 1980s sustained into the 1990s. The NAIRU also declined over this period by about 1-2 percentage points according to the estimates of Stiglitz (1997, p6), Gordon (1997), Staiger, Stock and Watson (1996), FRB/US (1999) and the OECD (1996). However, this "coincidence" does not drive the results; it is what would have been expected on the basis of past relationships. The estimated coefficient is similar using data only up to 1980. The low frequency variations in the data strengthen my results, but they are not necessary for them.

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[5] 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.

The stability of the minimum wage coefficient before and after 1980 also suggests that the results cannot be accounted for by compositional changes in the wage data. If employment of unskilled labor grew more slowly than employment of skilled labor when the minimum wage was high, then the average wage would rise. In principle, compositional changes like these could account for the correlation seen in the data up to 1980. However, they would not account for movements since 1980 in the Employment Cost Index, which uses fixed occupational weights. (Nor would they account for the evidence of Section 1.3). In any case, these effects do not seem to be important. Employment of 16-19 year olds has grown more slowly than that of adults over the last two decades despite the low level of the minimum wage.

1.4.2 Sensitivity analysis

To guard against misspecification and omitted variable bias, and as a guide to future research, Table 1-3 shows some tests of exclusion and other restrictions imposed in estimating my wage equation. The table also shows the effect of excluded variables on the coefficient on the level of the minimum wage. Typically new variables are included with both the current and one lagged value, to allow for both level and change effects. Where new variables are collinear with other regressors, I constrain the other regressors to equal their baseline estimates.

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 1953:1 to 1996:4, corresponding to the data availability of some of the excluded variables, the coefficient (x100) is 0.59, with standard error 0.16.

Table 1-3: 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 prod’y	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*

Pricewedge	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	
Unemployment above NAIRU	61%	0.74	0.21	
reciprocal of unemployment rate	29%	0.89	0.17	
logarithm of unemployment rate	23%	0.90	0.17	
III. Rejected restrictions				
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 1-3:

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: 1976:3 to 1997:1
c) as with mean duration, this was lagged twice, matching the lag on unemployment. The proportion unemployed over 15 and 27 weeks gave 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. See Section 2.3.1.
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 increase with the inflation rate. See Section 2.3.3
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 attaches to 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. The sensitivity of the unemployment effect will need to be taken into account when research proceeds beyond the question of “whether” to the question of “how much”. But we are not at that stage yet.

I discuss several of the restrictions in the table in later chapters. In Section 3.6, I argue that the insignificance of other elements of the social safety net (the first five rows) suggests that the minimum wage is unlikely to simply serve as a proxy for deeper social forces, such as readiness to intervene in markets or egalitarian attitudes. As I discuss in Section 2.7, the existence of a unique NAIRU is an 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. I discuss some more restrictions in Section 2.2.

1.5 Conclusion

A positive effect of the minimum wage on the NAIRU is evident in four data sets:

- wage growth from 1948 to the 1970s
- wage growth from the 1970s to 1998
- the behavior of prices
- international comparisons of changes in the NAIRU

Reflecting this, it is consistent with – and explains – time-varying estimates of the NAIRU.

Although these sources of information suggest slightly different point estimates, they are all consistent with a 10 per cent increase in the relative minimum wage raising the NAIRU by about half a percentage point. As I discuss in Chapters 3 and 4 this result is corroborated, in qualitative terms, by information on worker insecurity and is consistent with theoretical models of wage determination.

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 factors could also underlie 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 (Section 1.4.2) were unsuccessful.

The other sources of information can also be challenged. My price-price equation is sensitive to alternative specifications. My cross-country comparisons may not survive multivariate analysis. The evidence on worker insecurity is open to alternative interpretations.

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 1 -- 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 1-4: A Price Markup Equation

Regressor	**Coefficient**	**Std. 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	1955:1 1998:1
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 (1957:2 – 1996:1)	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 also the same, except that it is not averaged and lagged. 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 1.3.2). 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 markup, not an ongoing change in inflation. The long run effect of the level of the minimum wage on the NAIRU is not affected. In the notation of section 1.2.1, the derivative of the NAIRU with respect to elements of X – that is, δ/β – is independent of whether those elements are also contained in Z.

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 Gordon's finding that unit labor costs were "irrelevant for inflation".

When these variables are measured for the same sector, as above, strong relationships between them are evident. In a general specification that includes four lags of inflation, these extra terms are 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 3rd 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 may nevertheless be 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 might be proxied more accurately by lagged inflation than by unit labor costs for the business sector.

Appendix 2: Data

Table 1-5: Data Measures and Sources

Variable	Measure	Data source
Wages from 1980:1	Compensation of non-farm private industry workers	Employment Cost Index, BLS
Wages to 1980:1	Compensation per hour; non-farm business sector	unpublished BLS data, from "Productivity and Costs"
Prices from 1947:1	Chain-weighted price index for personal consumption expenditure	NIPA Table 7.1
Prices 1942 to 1947	Implicit price deflator for personal consumption expenditures at 1987 prices; interpolated from annual data	BEA (1993); Table 7.1 row16
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 below
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 1980:1).	Social security Administration (1998)
wage freeze	1 in 1971:4; -0.6 in 1972:1	FRB/US database (DWPC)
food and energy prices	Difference between deflators 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	deflator for agricultural output	FRB/US database (PH.GDP_D87YAGF)
energy prices	price of oil to 1956 then composite of prices for oil, coal and gas.	FRB/US database (PCENG, POIL)
price controls	1 from 1971:3 to 1974:1; -3.67 from 1974:2 to 1974:4	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 1973:1. For the measure used in the wage equation, I then take a 5-year back average of the quarterly change in trend productivity.

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. The average divergence between the series from 1950 to 1998 is 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, 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).

Figure 1.4 below shows the logarithm of the minimum wage, divided by average hourly compensation, and this series adjusted for coverage.

Figure 1.4: The relative minimum wage