Topics in Macroeconomics

Volume 4, *Issue* 1 | 2004 | *Article* 7

Do Minimum Wages Raise the NAIRU?

Peter Tulip*

*Federal Reserve Board, ptulip@frb.gov

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Abstract

Maybe a lot. The ratio of the minimum wage to the average wage enters a Phillips-curve equation with a coefficient that is highly significant, stable, and robust. One interpretation of these results is that the relative level of the minimum wage affects the Non-Accelerating Inflation Rate of Unemployment or NAIRU. My estimates are consistent with the reduction in the relative level of the minimum wage since 1980 lowering the U.S. NAIRU about $1\frac{1}{2}$ percentage points, while raising the NAIRU in continental Europe. However, other interpretations are also possible.

KEYWORDS: NAIRU, minimum wage

*Federal Reserve Board, Washington D.C. 20551. Email: ptulip@frb.gov. This paper is a substantially revised version of Tulip (2000b). I am grateful to Charles Fleischman, Flint Brayton, Dave Reifschneider, Albert Ando, Lawrence Klein, Dan Polsky, Frank Schorfheide, Clive Crook, David Romer and an anonymous referee for helpful comments and assistance. The data and programs used in this paper are available from the author on request and at http://www.petertulip.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 Alan S. Blinder (1997) and Edward M. Gramlich (1998) have used them to guide monetary policy. Joseph Stiglitz (1997) finds them useful for framing policy discussions. Professional forecasters use them for forecasting inflation (Federal Reserve Bank of Philadelphia, 2001). Richard Layard et al. (1991), the Organization for Economic Co-operation and Development (OECD, 1994) and Edmund S. Phelps (1994) use them as a measure of the permanent component of unemployment.

In this paper I investigate whether the NAIRU is affected by the level of the minimum wage (relative to the average wage). I find that it possibly is and that this effect may be important. For example, movements in the level of the minimum wage may account for much of the upward drift of the NAIRU in the United States over the 1960s and 1970s and its subsequent decline. They may also help explain why the NAIRU (and hence unemployment) has risen in continental Europe while falling in the United States.

If the effect of the minimum wage on the NAIRU is large, it will be relevant to macroeconomists. And even if it is small, it should be of interest to those assessing the minimum wage. This is because the effect on the NAIRU represents the *sustainable* unemployment arising from the policy. In contrast, a change in actual unemployment, unaccompanied by a change in the NAIRU, would result in accelerating or decelerating prices. Because this cannot be sustained, the actual rate of unemployment will converge to the Non-Accelerating Inflation Rate of Unemployment.

The change in the NAIRU is arguably a more useful measure of the consequences of the minimum wage than the conventional focus on changes in employment (surveyed by Charles Brown, 1999). 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, it is the subject of this paper. But it 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.

The idea that the minimum wage may increase the NAIRU is not new. Immediately after introducing the concept of a natural rate of unemployment, Milton Friedman (1968, p.9) suggested that it would be affected by the minimum wage. However, this conjecture has not been verified and the idea is barely mentioned in surveys of the minimum wage (Brown, 1999) or of the NAIRU (for example, Layard et al., 1991, Katz and Krueger, 1999). The OECD (1997a) and R. Jackman and C. Leroy (1996) have found the minimum wage to have large effects on the NAIRU in France, but the results do not appear to be strong and details have not been published. Other researchers have failed to find significant effects. Jackman et al. (1996, n.2) 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 Douglas Staiger et al. (1997) find that the minimum wage is insignificant in explaining the acceleration of prices in the United States, but do not report estimates.

My research differs from these earlier efforts in that I focus on wage determination in the United States, where the data permit more powerful tests. Specifically, I find a significant effect of the ratio of the minimum wage to the average wage in an equation explaining the growth in nominal wages. Under certain conditions, which I explore, this correlation can be interpreted as a causal effect. If this effect flows on to prices – as it appears to – then extra unemployment would be required to offset the inflationary pressure, so the NAIRU is higher.

My results are an extension of the large literature that finds the minimum wage places upward pressure on other wages – so-called "ripple" effects. Examples include Gramlich (1976), Jean Baldwin Grossman (1983), William E. Spriggs and Bruce W. Klein (1994), and David Card and Alan B. Krueger (1995). Whereas these studies focused on immediate effects, I find persistent, and hence much larger, effects. My results are more closely related to those of F. Gerard Adams (1989) who found that the gap between the average and minimum wages is significant in a wage equation, though Adams did not explore this finding or its implications for the NAIRU. My results are also closely related to the literature on the correlation between unemployment and wage compression in cross-country comparisons. Section 5A gives citations and suggests that this literature corroborates and complements my results.

The plan of the paper is as follows. Section 2 explains how and why estimates of the NAIRU can be derived from a wage equation. Section 3 presents a wage equation in which the minimum wage is highly significant. Section 4 notes some doubts: for example, the correlation might be accidental or it might not reflect a causal relationship. The following two sections help to allay some (though not all) of these concerns. Section 5 shows that the results also help explain international comparisons and the behavior of prices. Section 6 shows that my results are not unduly sensitive to variations in sample period or specification. Section 7 concludes.

It should be noted that many economists are skeptical of NAIRU estimates. This is a methodological debate that this paper does not address, beyond some brief comments in Section 4. Uses of the NAIRU, such as those noted in the first paragraph, seem sufficiently influential that investigation of its determinants seems worthwhile – and this might most effectively be done within the framework of the existing literature. Of course, those who are skeptical of this literature will regard this limitation as serious.

The focus of the paper is empirical. It seems useful to determine whether minimum wages have an important effect on the NAIRU before asking how or why. 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 unchanged inflation.

The minimum wage can also increase the NAIRU if it causes other nominal wages to gradually increase, for 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 Jeremy I. Bulow and Lawrence H. Summers (1986). These effects can take some subtle forms. For example, Card and Krueger (1995, p.163) find that the minimum wage sometimes affects starting wages but not the slope of wage-tenure profiles. Then an increase in the minimum wage would slowly flow to other 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 is supported by – and helps to explain – the widespread view that the reduction in the NAIRU in the 1990s was due to worker insecurity.

In Tulip (2000a, Ch.3) 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 other workers – and from them, to still more workers. There are indications that these wage-wage interactions may be important, though the evidence is not strong. But these matters of interpretation lie outside the scope of this paper.

2. DERIVING THE NAIRU FROM A WAGE EQUATION

Inferences about the NAIRU are typically obtained from estimating a "backward-looking Phillips curve," in which the change in price or wage inflation is regressed against an activity variable, such as the rate of unemployment. The specifications of these regressions are chosen with a view to fitting the data rather than being derived from assumptions about preferences. This approach is controversial and I discuss some doubts about its validity in Section 4.

Within the NAIRU literature there are several variations. Perhaps the most common approach is estimation of a "price-price Phillips curve" in which inflation is regressed on lagged inflation and other variables. Examples include Robert J. Gordon (1997, 1998) and Flint Brayton et al. (1999). Alternatively, one can include the change in inflation as a regressor in an equation that explains unemployment – "inverting the Phillips curve". Examples include Layard et al. (1991, p.55), Phelps (1994, p.314) and Giuseppe Bertola et al. (2001). A significant effect of the level of the minimum wage on the NAIRU can be found using these methods.[2] However, I follow a third approach, decomposing inflation into separate price and wage equations, which I suspect provides clearer, more reliable estimates.

Derivations of the NAIRU from interacting wage and price equations can be found in macroeconomic textbooks and elsewhere. Olivier Blanchard and Lawrence F. Katz (1997, p.60) and Katz and Krueger (1999, p.15) provide simple and accessible presentations. Following Albert Ando and Flint Brayton (1995), my framework differs slightly from the textbook approach in that I impose cointegration between prices and unit labor costs and I distinguish between consumer and product prices. In this framework, prices appear to be a stable markup on unit labor costs. Accordingly, it then seems appropriate to focus on the determinants of nominal wage growth.

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. To simplify the exposition, I ignore most lags including that of the dependent variable. Measuring lower case variables in logarithms and letting Δ represent differences, the wage equation can be written:

$$\Delta w = \Delta p_{c(-1)} + \Delta prod + \beta U + \delta X. \tag{1}$$

The estimated equation is intended to capture the main empirical influences governing wage growth – though how well it does so clearly depends on lag

[2] For a price-price Phillips curve, see Tulip (2000b). For an inverted Phillips curve, see the discussion of Bertola et al. in Section 5A.

structures and on how X is constituted. I discuss its specification in the following section. Consistent with the Phillips curve literature, 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 remain unclear, despite active research.

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

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

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. Examples include Ando and Brayton (1995, p.289), Brayton et al. (1999, Table A5) and Tulip (2000b, Appendix 2).[3] The key result for present purposes is that the error-correction coefficient, λ, is significantly less than zero. This means the equation determines a relationship between the level of prices and the level of unit labor costs. Specifically, there is a target level of prices, p_p*, given as:

$$(3) \qquad p_p* = w - prod - (\gamma U + Z) / \lambda .$$

As can be seen by substituting (3) into (2), prices will remain near this level. Otherwise, say if $p < p*$, then prices would rise faster than unit labor costs, bringing the markup back to target.

In the modern U.S. economy, U and Z are approximately stationary (more precisely, their trending components are small), as are the inflation rate, the relative minimum wage and many other variables that could be included in the specification. So differencing (3) and setting ΔU and ΔZ to their mean values of zero, gives a version of equation (2) that explains long-run changes in prices:

$$(4) \qquad \Delta p_p = \Delta w - \Delta prod.$$

To put this another way, prices and underlying unit labor costs appear to be cointegrated (equivalently, the labor share is stationary) and this arises through prices error-correcting to costs.[4] Assuming that trend productivity is exogenous, that means the long-run inflation rate is explained by the determinants of wage growth.

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[3] Estimates from an updated and revised equation are available on request.

[4] Why not show this directly with, for example, a Dickey-Fuller test? Partly to show that it is prices that adjust and partly because, as Kremers et al. (1992) argue, t-tests of the error-correction coefficient λ provide a more powerful test of cointegration.

Assuming that prices follow wages differs from the popular alternative of explaining inflation through a "price-price Phillips curve." This latter 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 not strongly correlated across sectors, hence the apparent finding that unit labor costs are "irrelevant for inflation." However, when these variables are measured for the same sector, then unit labor costs become highly significant in both statistical and economic terms.

An appropriate measure of prices for assessing the effect of labor costs is the price of business sector output ("product prices"), which is the broadest sector for which prices, wages and productivity can be measured on a consistent and reliable basis. However the measure that affects wages is consumer prices. (If a weighted average of the two price series is included in the wage equation, the estimated weight on product prices is approximately zero.) An equation links these series:

$$\Delta p_c = \Delta p_p + wedge. \tag{5}$$

The difference between consumer prices and product prices, which I call *wedge,* is a composite of many influences, which I take as given. In the short term, these include fluctuations in farm prices and the external terms of trade. In the longer term, product prices have risen less than consumer prices, reflecting faster technological change in the production of investment goods relative to consumption goods.

A "short-run NAIRU" could be estimated by numerically solving equations (1), (2) and (5) for the unemployment rate at which inflation is stable.[5] I focus instead on a long-run measure of the NAIRU, derived by using equation (4) instead of (2). This is simpler and seems more relevant to policy analysis. Substituting (1) into (4), then (4) into (5) gives a reduced form for inflation that applies once price margins have returned to their long-run levels:

$$\Delta p_c = \Delta p_{c(-1)} + \beta U + \delta X + wedge. \tag{6}$$

Setting $\Delta p_c = \Delta p_{c(-1)}$ and solving for the unemployment rate gives:

$$NAIRU = -[\delta X + wedge] / \beta. \tag{7}$$

[5] This measure of the NAIRU reflects shocks to both wage and price equations (that is, the *X* and *Z* variables) being offset by the effect of unemployment on both wages and price margins (the coefficients β and γ). Such estimates are useful for some purposes – for example, assessing temporary influences or comparison with price-price Phillips curves. However, a limitation of this measure (and, for that matter, of those based on price-price Phillips curves) is that it does not abstract from transient changes in the markup. Shocks that are unsustainable do not need to be offset by unemployment.

This represents the unemployment rate required to eventually stabilize inflation, when allowance is made for the tendency of fluctuations in price margins to disappear of their own accord. 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 than unit labor costs. This causes a temporary deviation of price margins from the long-run level given by equation (3), which is gradually eliminated by the error-correction term.

A few other simplifications may deserve noting. In the long run the NAIRU is assumed to be independent of productivity, which eventually is fully reflected in nominal wage growth, leaving unit labor costs and hence prices unaffected.[6] Vacancies are not included within X, so measures of "frictional unemployment" based on shifts in the Beveridge curve do not directly affect the NAIRU. The lagged wage share is also excluded, which means that, in the wage-unemployment diagrams of Layard et al. (1991, p.14), Phelps (1994) or Blanchard and Katz (1997, p.55) my "wage setting" or "labor supply" schedule is vertical. Allowing for these complications would involve modifications to the algebra above. However, as shown in Table 2 below, they do not seem to have been important in the United States – though they may matter elsewhere.

Equation (7) provides a framework for estimating contributions to the NAIRU. Specifically, the vector X comprises variables that have an important effect on wages and the parameters δ and β can be estimated from an equation like *(1)*. To obtain estimates of the level of the NAIRU (a distinct question), some average of *wedge* can be included.

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 a small negative value of β is well established (see for example Layard et al., 1991, p.199, or the many references cited by Robert W. Rich and Donald Rissmiller, 2001), I am especially interested in estimates of δ^* – the contribution of the minimum wage to nominal wage growth.

[6] Because trend productivity affects wages with a lag, changes in the trend *temporarily* affect unit labor costs and the "short-run NAIRU" discussed in footnote 5. Coupled with estimates for equation (2) (not shown), my wage equation implies that the deceleration in productivity boosted the short-run NAIRU by an average of three-quarters of a percentage point from 1973 to 1979. An acceleration of productivity lowered it half a percentage point from 1995 to 2001.

3. ESTIMATES

Table 1 presents an econometric equation explaining nominal wage growth in the United States – a fleshed out version of equation (1). The dependent variable is the quarterly percentage change (more precisely, the log difference times 100) in compensation per hour in the non-farm business sector. Explanatory variables include lagged wage growth, productivity growth, inflation, unemployment, the relative minimum wage, payroll taxes and the 1971 wage freeze.

Following convention, I constrain the sum of coefficients on inflation and lagged wages to equal 1 so that there is no long-run tradeoff between inflation and unemployment. I also constrain the sum of coefficients on productivity and lagged wages to equal 1, so that real wages and productivity grow at the same long-run rate. Neither restriction is rejected by the data (see Table 2 below). Smoothness restrictions, in the form of moving averages, are imposed on lags for simplicity. Lag structures are chosen by searching over many alternatives and selecting that which minimizes the Schwarz criterion subject to the homogeneity and smoothness restrictions. This results in lags of up to 5 years on both inflation and trend productivity growth. Although these long lags may seem surprising, they are common in the Phillips-curve literature – see for example, Gordon (1998) or Brayton et al. (1999).

The specification closely resembles other recently published wage equations at a conceptual level. See, for example, those listed by Rich and Rissmiller (2001). The main difference is my inclusion of the *level* of the minimum wage, though there are also differences of detail. Of course, details can matter. But rather than debate issues of specification, I compare variations in Section 6.B and show that my results are not unduly sensitive to these.

Because there are hints of heteroskedasticity in some variations I consider, I report heteroskedasticity-robust "jackknife" standard errors of Russell Davidson and James G. MacKinnon (1993, p.554) in column 4, and throughout the paper, unless otherwise stated. (The conventional standard error of the coefficient on the level of the minimum wage is about the same.) I do not report significance levels, but all coefficients have t-statistics greater than 3.

The poor performance of some diagnostic tests suggests that the specification could be improved. In particular, the payroll tax coefficient is unstable (a functional form problem, I suspect) and residuals are not normal (reflecting a blip in 2000:q1). Simple controls for these problems do not seem to affect other coefficients much. Furthermore, my treatment of some effects, in particular trend productivity, is simple. But these issues are left for future research.

Table 1: Wage Equation

Dependent variable: (Δw) compensation per hour, nonfarm business sector, log difference, multiplied by 100

Regressor	Specification [a]	Coefficient	Standard error	Stability p-value [b]
Lagged wages	$\Delta^2 w_{(-1)}$	.28	.088	.15
Previous quarter's inflation	$\Delta p_{c(-1)}$	.34	.073	.45
Preceding 5 years' inflation	$\Delta^{19} p_{c(-2)}$	.38		
Trend productivity growth	$\Delta^{20} prod$	.72		
Minimum wage (level)	$(M_{(-1)} / AHE_{(-1)})$ x $COVERAGE_{(-1)}$	2.1	.62	.55
Minimum wage (change)	$\Delta m - \Delta^4 w_{(-1)}$	.042	.005	.99
Unemployment rate	$UDEM$	-.103	.026	.21
Payroll tax	see data appendix	1.0	.20	< .001
1971 wage freeze	see data appendix	-.98	.16	
Constant		.45	.15	.34

Diagnostics and other statistics [c]

Sample	1948:q1 to 2003:q1 (221 observations)
Standard error	.4066
R-Squared	.663
RESET test of functional form	$p = .44$
Breusch-Godfrey test for up to 4th order serial correlation	$p = .96$
White's test for heteroskedasticity/misspecification	$p = .29$
Jarque-Bera test for normality	$p < 0.001$
Overall stability (breakpoints from 1951:q1 to 1999:q4)	$p = .31$

Notes:

(a) Lower case variables are in logarithms. Δ^n represents the average change over n quarters multiplied by 100 (approximately the average percent change). Detailed variable definitions are in the Appendix. In brief, *prod* is trend labor productivity, with kinks at 1973:q1 and 1995:q1; *UDEM* is the unemployment rate in percentage terms, with fixed demographic weights; *M* is the average minimum wage across states; *AHE* is average hourly earnings; and *COVERAGE* is the proportion of private sector employees covered by the Federal minimum wage legislation. (*M / AHE)* x *COVERAGE* is the Kaitz index, the usual measure of the minimum wage in empirical studies (Brown, 1999, p.2114). The Kaitz index is entered as the difference from its 2003:q1 value.

(b) Bruce E. Hansen's (1997) *p*-values for the Andrews-Ploberger Exp-F statistic for all breakpoints between 1951:q1 and 1999:q4, holding other coefficients constant.

(c) Diagnostics are reported as *p*-values, with low values implying rejection of the "classical" assumptions. RESET represents a t-test on the addition of the squared fitted values to the equation. Overall stability represents an Exp-F test, as in (b), for simultaneous breaks in all coefficients except the wage freeze and payroll tax, which are constrained to their full-sample estimates. Other statistics are calculated by Eviews 4.1 using F-statistic versions assuming homoskedasticity.

The coefficient on the *change* in the minimum wage means that a 10 percent increase in the relative minimum wage would immediately boost aggregate nominal wages by about 0.4 percent. They would then rise further owing to wage-wage and wage-price interactions. These effects would substantially boost estimates of the "short-run NAIRU." But in the long run, these effects fade to zero – assuming the relative minimum wage does not grow indefinitely.

In addition to these temporary effects, the higher *level* of the minimum wage, relative to the average wage, would continue to raise aggregate wages, for a given level of unemployment. A 10 percent increase in the relative level of the minimum wage, from its level at the end of the sample, would boost aggregate wage growth by 0.07 percent a quarter.[7] This continued effect is small in some respects. It is less than rounding errors in the data and substantially less than the standard error of my wage equation. So powerful tests – that is, good controls and a long data set – are necessary to see it. But with these the effect is precisely estimated and clearly discernible. The coefficient is 3.4 times as large as its estimated standard error. And, as I show in Section 6, the estimate is fairly robust to variations in sample period and specification.

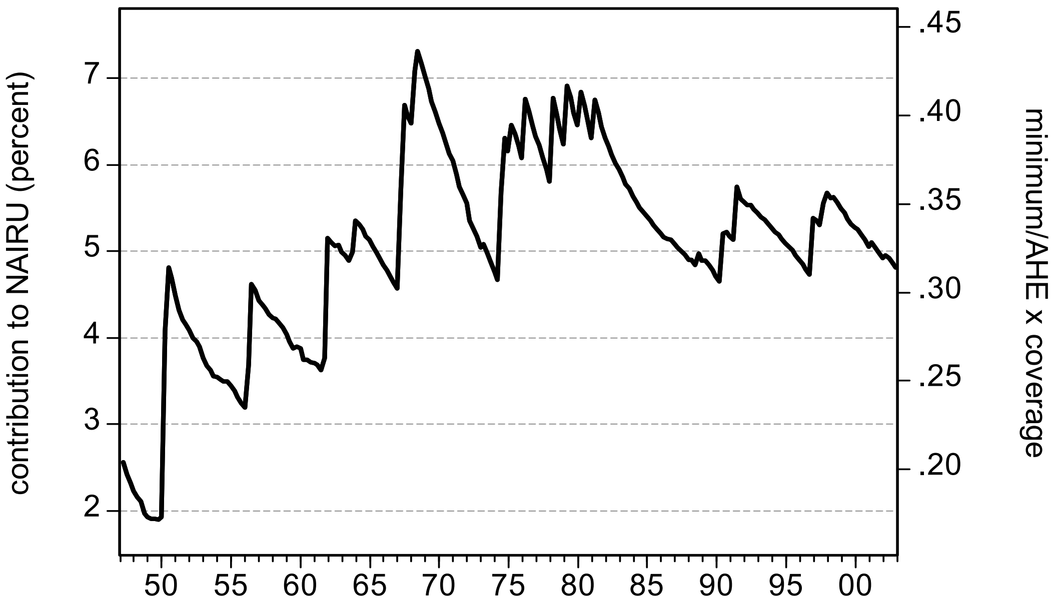
Although the effect on nominal wages may seem small, it is large relative to the effect of unemployment on wages, which is the metric that matters for policy. This is estimated by the coefficient on unemployment in Table 1, -0.103, an estimate that is similar to those of other researchers. For example, Gordon's (1998, Table 3) ECI equation has an unemployment coefficient of -0.12, when divided by four to be on a comparable quarterly basis. Rather than taking unemployment as given (as in the previous paragraph), which would imply ever-increasing inflation, it seems more relevant to assume that higher unemployment offsets the rising wages. To offset the 0.07 percent wage growth discussed above would require an extra 0.7 percentage points (=.07/0.1) of unemployment. That is, a 10 percent increase in the minimum wage raises the NAIRU by a bit more than half a percentage point.

Chart 1 shows historical estimates of the (demographically adjusted) NAIRU implied by the level of the minimum wage. I include relevant constants (the intercept from the wage equation and the sample average of the wedge), so the

[7] With my specification this effect increases with the relative level of the minimum wage. In 2003:q1 the average minimum wage was $5.47 (slightly above the Federal minimum of $5.15), coverage was 86 percent and average hourly earnings were $15.05. So the relative minimum wage, or Kaitz index, was 0.31 (= 5.47 x 0.86/15.05). The coefficient on this term is 2.1. So a 10 percent increase in the Kaitz index raises wages by 0.066 percent (= 2.1 x 0.1 x 0.31). In Tulip (2000b) I used a logarithmic specification, which is simpler to interpret and fits the data better. But the linear specification seems more plausible outside my data range and consistent with the larger effects apparent in international comparisons. The appropriate functional form is an issue on which more work could be done.

level of the series can be interpreted as the NAIRU at a particular point in time. For reference, the right hand axis shows units for the Kaitz index, that is, the level of the minimum wage, divided by average hourly earnings and multiplied by coverage.

Chart 1: The NAIRU[8]
Contributions from the level of the minimum wage and constants



The chart indicates that when the minimum wage was low, as in the 1950s, the NAIRU fluctuated around 4 percent. Movements in the minimum wage can then account for a rise in the NAIRU to around 6½ percent in the 1970s and then a reduction to 5 percent in the 1990s. To be more precise, the coverage-adjusted minimum wage declined from 40 percent of average hourly earnings in the late

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[8] The plotted series is obtained by substituting coefficients from Table 1 into equation (7):

$$\text{(8)} \qquad \text{NAIRU} = \underset{[0.4]}{4.8} + \underset{[5.7]}{20.5}\ \text{minimum wage}$$

where the constant equals the sum of the wage equation intercept and the average wedge, both divided by the unemployment coefficient ((0.45 + 0.04)/0.103 = 4.8). The minimum wage is measured as the deviation from its latest value so the constant is interpretable as the NAIRU in 2003:q1. Standard errors, estimated by nonlinear least squares, are in brackets. The (relatively small) standard error for the constant assumes the wedge is fixed.

1970s to 33 percent of average hourly earnings over the 1990s. This reduced the NAIRU by about 1½ percentage points ((.40-.33) x 2.1/0.103 = 1.4).

The size of this effect may seem surprising. In part, this is because it is often not appreciated how flat the short-run Phillips curve is. As Gordon (1997, p.27) and others have noted, it actually takes a lot of unemployment to offset small inflationary shocks. It may also seem odd that such a large effect has so far gone unnoticed. However, that is only the case in the United States, where, until recently, the NAIRU was conventionally modeled as constant. In other countries similar effects have attracted widespread comment, as Section 5A documents. And the large effect may be perceived to be at variance with other research. In particular, Card and Krueger found negligible effects of the minimum wage on the demand for labor. However, this (controversial) result is not necessarily inconsistent with mine. The minimum wage might affect the NAIRU by some means other than labor demand – for example, wage-wage interactions or labor supply. It should be acknowledged, however, that microeconomic evidence of these other channels of influence is not strong.

4. DOUBTS AND RESERVATIONS

It is not clear how much confidence one should place in a t-statistic of 3.4, as I estimate for the minimum wage. If the test was independent of other research, and if the coefficient had a normal or t-distribution, then its *p*-value would be about 0.1 percent. However, this understates the likelihood of obtaining my results for several reasons.

One minor concern is the non-standard distribution of the coefficient. Because the relative minimum wage is the ratio of two integrated variables, its coefficient may have fatter tails than a t-distribution (see Kremers et al., 1992). And, because the regressor includes (a transformation of) the lagged level of the dependent variable, its coefficient, as with error-correction coefficients in general, will tend to be biased away from zero. However, a simple Monte-Carlo exercise suggests that these complications seem unlikely to be important for my full-sample estimates.[9]

[9] I construct two independent random walks each of 221 observations with normally distributed disturbances with variances equal to those of first differences in the logarithms of the aggregate wage and the coverage-adjusted minimum wage. In artificial regressions of the change in the first of these series (representing the change in wages) on the lagged difference between them (representing the relative minimum, under the null hypothesis), the average least-squares coefficient (times 100) is not zero but 0.05. This bias seems small when compared with my estimate (the elasticity at the last observation, or from a logarithmic transformation) of 0.7 – though it may justify downward rounding. Whereas a conventional one-tailed t-test would assume that the

A bigger concern is that my equation is a product of an industry-wide specification search. Because surprising results are more likely to be reported, published *p*-values understate the likelihood of incorrectly retaining unimportant variables (Frank Denton, 1985). To guard against this, and to explore further implications, it is informative to check other kinds of evidence. One gauge of the plausibility of my estimates is the extent to which they are consistent with other data. I explore this in the following section.

Yet even if other data sets point to similar relationships – as they do – these correlations need not imply that the minimum wage "causes" wage growth. For example, one might worry that the causation flows in the opposite direction. As there is no serial correlation in my residuals, and the minimum wage is lagged, this would mean the future "determining" the past. This can occur when policy is based on forecasts (as, for example, with monetary policy), but that does not seem to be important with the minimum wage. Congress sets the relative level of the minimum wage for many years in advance (subject to some small uncertainty about the level of the average wage). In doing so, it shows little interest in, or knowledge of, the unexplained change in inflation that is likely to occur over that period. So residuals from the wage equation seem unlikely to influence the lagged level of the minimum wage.

Of course, Congress may be reacting to lagged inflation, unemployment, or some other variable that also determines subsequent wage growth. However, the nature of multiple regression is that the significance of the minimum wage represents the information it provides beyond that contained in other regressors. Of more concern is the possibility that some omitted variable is correlated with both the minimum wage and the NAIRU. In this case the minimum wage would still be informative for forecasting, but the correlation would become unstable if used for policy purposes. To help assess this, in Section 6 I test the stability of the minimum wage coefficient and for the presence of omitted variables. One can never rule out the possibility of omitted variable bias, but these tests help to allay many concerns. They provide some basis for the assumption that the significant coefficient of the minimum wage reflects a causal effect of the minimum wage on the NAIRU. But further research on the validity of this assumption would be helpful.

Even if there were a highly significant, stable and robust coefficient, many economists would regard this evidence as unpersuasive because the backward-looking Phillips curve is not derived from formal theory. They argue that we

probability of obtaining a t-statistic of 3.4 or greater is 0.04 percent, the probability based on my artificial data is about twice as high, 0.09 percent, but still tiny. Of course, my assumptions about log transformations, normal distributions, the absence of impact effects and the absence of other regressors are simplifications. But the small bias arising from the presence of the lagged level of the dependent variable does not suggest that a more complicated analysis would be informative.

cannot rely on relationships that we do not understand. Specifically, the conditions under which estimated correlations will remain stable are unclear. It may be that these reduced form relationships reflect expectations regarding policy or responses to a particular constellation of shocks and hence they will break down when the policy or the nature of the shocks changes.

This concern applies to the NAIRU literature in general. Few, if any, of the NAIRU estimates used by policy-makers or forecasters are based on formal microeconomic foundations. Many papers have been written on the utility of this approach (the introduction cites a few), and I do not wish to add to them here. In brief, my position is that the absence of theoretical underpinnings – and microeconomic evidence, for that matter – reduces confidence in the reliability of my results. Conclusions drawn on the basis of incomplete information are tentative. However, pending development of a reasonable theory of the Phillips curve, policy needs to be made on the basis of available information. If empirical relationships have been stable over extended periods of time and across substantial changes in policy, then one has reason for assuming that these relationships may continue to hold. If the relationships are consistent with other information, particularly from different economies, then confidence in this assumption is increased. In that sense, the NAIRU literature and the approach it follows are informative.

5. FURTHER IMPLICATIONS

5.A. International comparisons

Variations in the NAIRU and the minimum wage are often larger within and between other countries than within the United States. So international comparisons should provide corroboration and interesting extensions. This section notes some direct then indirect comparisons. The indirect comparisons are more powerful, but their relevance depends on stronger assumptions.

A few other large economies have national statutory minima, as in the United States. OECD estimates of the NAIRU for these countries seem to be positively, though imperfectly, correlated with trends in the relative minimum wage.[10] In France, both the relative minimum wage and the NAIRU have risen by large amounts over the last few decades. In Portugal, both have fallen. In the Netherlands, both rose and then fell, though the NAIRU is estimated to have

[10] For estimates of the minimum wage from 1970, see OECD (1998, Chart 2.2). For the NAIRU, see OECD (1996b). OECD country studies provide earlier estimates and further information.

peaked after the minimum wage. Spain is the major exception: although the Spanish minimum wage has fallen, its NAIRU has risen sharply.

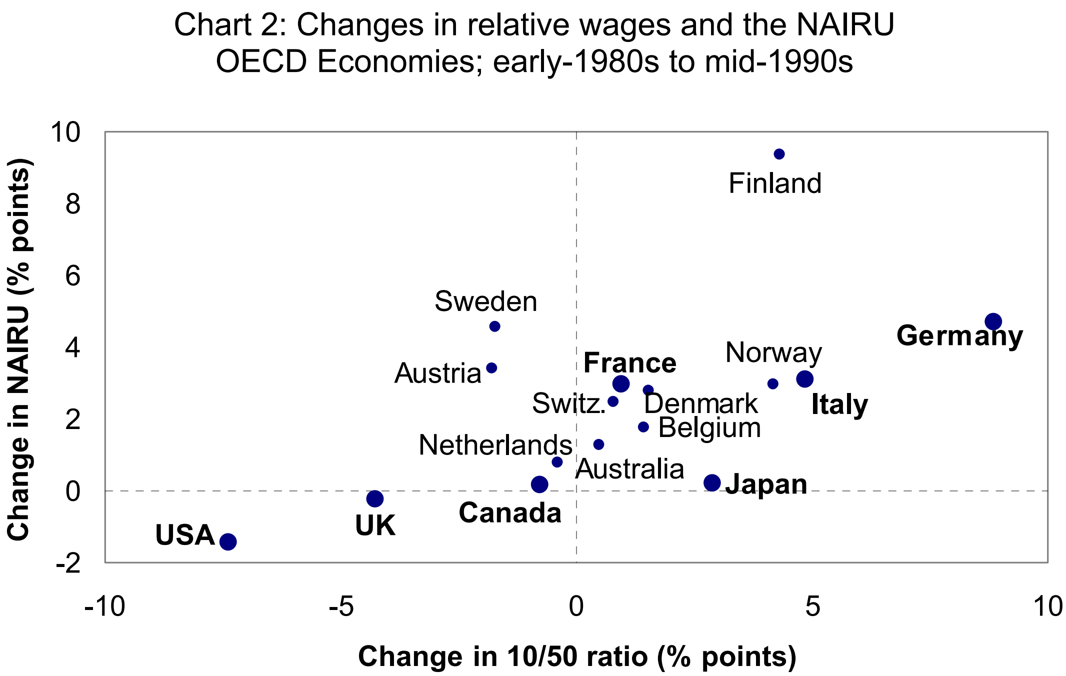
Multivariate analyses have been conducted for France. The OECD (1997a) and Jackman and Leroy (1996) find the effect of the minimum wage on the French NAIRU to be statistically significant. The OECD (1997b) concludes from these results that a 10 percent increase in the French minimum wage, relative to the average wage, would increase the NAIRU by 0.9 percentage points. Allowing for the higher relative level of the French minimum wage, this is in line with my estimates for the United States.

Other governments support low wages in other ways, for example by extending collective bargaining agreements. If we assume that these interventions have similar effects to the minimum wage, then we greatly extend the scope for comparisons. And if we assume that these attempts to boost low wages can be proxied by the ratio of the 10th percentile of the wage distribution to the median, referred to as the "10/50 ratio", then they can be quantified and compared. In U.S. and French data the 10/50 ratio is strongly correlated with the ratio of the coverage-adjusted minimum wage to the mean – as might be expected given the proximity of the minimum and mean wage to the 10th and 50th percentiles of the wage distribution.

Chart 2 plots changes in the NAIRU between 1980 and 1995 against changes in the 10/50 ratio over similar periods. The G7 are in bold. Data on the 10/50 ratio are compiled by the OECD (1996a).[11] For estimates of changes in the NAIRU, I use the "NAWRU" series compiled by the OECD Secretariat (1996b, Table 1). These are based on the bivariate relationship between wage growth and unemployment, adjusted to reflect additional research, where available.[12]

[11] Table 3.1. This provides sparse estimates from 1979 to 1995. I use the earliest and latest observation for each country, unless 1979 data are available, in which case I use the 1979-81 average. Data typically cover adult full-time employees. For the United States, I average the estimates for men and women.

[12] Earlier estimates have been published and explained by Jorgen Elmeskov (1993). Applications and discussions of the series can be found in Laurence Ball (1997, 1999) and numerous OECD reports on its Jobs Study Strategy.



Again, there is a noticeable correlation between relative wages and the NAIRU. The relationship is quantitatively similar to the U.S. and French results. 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 1½ percentage points higher in the mid 1990s. Chart 2 suggests a similar effect, perhaps slightly stronger.

The strength of the relationship in Chart 2 is sensitive to the weight given to the United States and to variations in measurement.[13] So it may be more interesting as illustration than as support. However, multivariate analyses provide stronger results. In particular, Bertola et al. (2001) regress international variations in unemployment on the male 10/50 ratio and numerous other variables. Among these is the change in inflation, meaning that movements along the Phillips curve (that is co-movements of inflation and unemployment) are controlled for and hence that other changes in unemployment can be interpreted as shifts of the NAIRU. Bertola et al. do not report coefficients, but they do note (p.195) that the 10/50

[13] In particular, the relationship is weak if the NAIRU is measured with the OECD's new price-based Kalman filter estimates, kindly provided to me by David Turner. Although the OECD prefers its new estimates, the earlier wage-based series is widely known, seems more relevant to my focus on wage determination and gives estimates for the United States that seem more consistent with other estimates of the time-varying NAIRU.

ratio is appropriately signed and statistically significant – much more strongly so than variables reflecting macroeconomic shocks.

If we assume that movements in the NAIRU and in actual unemployment are similar, as seems likely over longer time periods, then even stronger results are available. Numerous observers have noticed correlations similar to that shown in Chart 2, but using actual unemployment (or job loss) rather than the NAIRU. This "tradeoff" between unemployment and inequality has been documented for different samples, periods and measures and many papers have attempted to explain it. References include Paul Krugman (1994), the OECD Jobs Study (1994), Richard B. Freeman (1995), Giuseppe Bertola and Andrea Ichino (1995), Phelps (1997), Rebecca M. Blank (1997), Dale T. Mortensen and Christopher A. Pissarides (1999) and Bertola et al. (2001). The last of these appears to be the most thorough; the authors describe the partial correlation between the 10/50 ratio and unemployment as "very strong" (p.170) and relatively robust. Given that the correlation is typically observed in low-frequency data, it seems unlikely to be a cyclical phenomenon but rather to imply a correlation between the 10/50 ratio and the NAIRU.[14]

I conclude from the above discussion that if some strong but plausible assumptions are made, then international comparisons strongly corroborate the results of the previous section. That is, when wages at the bottom of the distribution are compressed, the NAIRU usually increases. Furthermore, the wide variety of policies across countries suggests that this correlation is not the result of one particular set of institutions or rules.

Implications also flow in the other direction. That is, the U.S. experience suggests that a large part of the increase in unemployment in Europe might be due to increases in the relative wages of low-paid workers. The U.S. time series not only provides evidence in support of the international tradeoff between unemployment and inequality, but it helps to interpret this correlation. As noted in Section 4, the U.S. minimum wage is predetermined many years in advance as, to a substantial extent, is the Kaitz index. So these variables are more easily assumed to be exogenous than the measures of inequality used in the international literature. The arguably exogenous nature of the minimum wage, and the similarity between the correlations in the U.S. time series and the international cross-section, suggest that the latter may reflect a causal effect of inequality on unemployment. This interpretation is simpler than the many attempts to find some third variable

[14] One caveat is that international variations in employment (and, to a less clear extent, unemployment) do not seem to be concentrated among the low-skilled, as might be expected if the correlation reflected shifts in the demand for labor. For evidence see Krueger and Jorn-Steffen Pischke (1997) and references cited therein. As with the U.S. studies noted earlier, evidence of weak demand effects might mean that some other mechanism, such as labor supply or wage-wage interactions, might be at work.

underlying the correlation. In any case, such attempts have not been clearly successful (see, for example, the discussion following Bertola and Ichino, 1995), and they seem to ignore the strong influence of governments in setting wages at the bottom of the distribution.

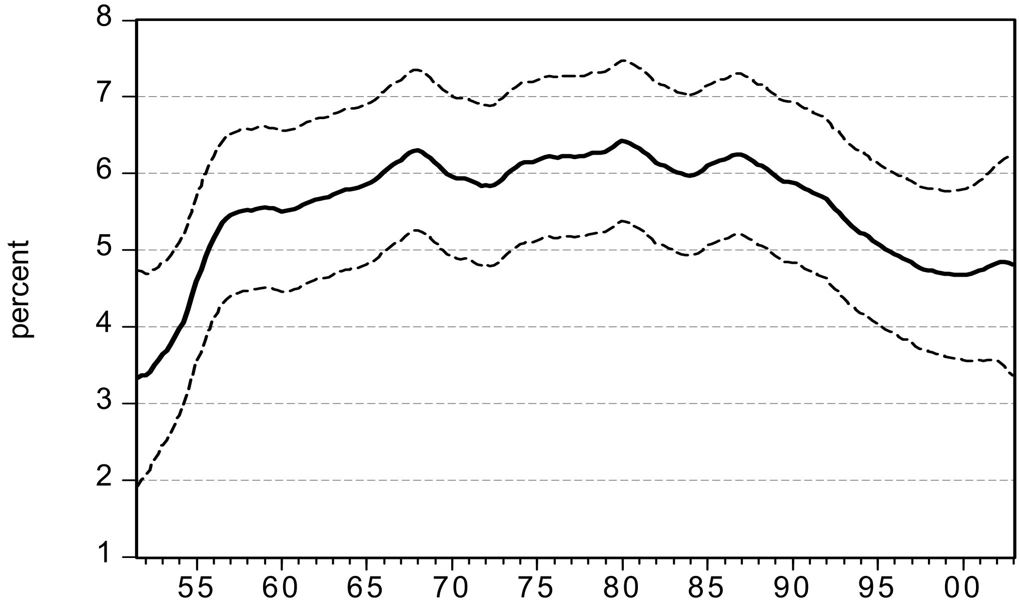
5.B. Consumer prices

Estimated price equations, such as those referred to in Section 2, find that labor costs flow into prices with a long, gradual lag. So transitory shocks to wages would not be evident in high-frequency fluctuations in prices. However, one would expect large persistent shocks, such as trends in the minimum wage, to be evident in longer term price movements. And so they are.

Chart 3 shows a two-sided Kalman filter estimate of a time-varying NAIRU. The estimates come from an equation explaining the change in consumer prices as a function of the demographically adjusted unemployment rate, three years of lagged inflation, the relative prices of food, energy and imports and an incomes policy dummy. The time-varying NAIRU follows a random walk, with variance estimated by the procedure of James H. Stock and Mark W. Watson (1998). The specification and results are similar to others (in particular the "PCEX" equation of Brayton et al.; see also the equations of Gordon, 1997, 1998). The main difference is that I extend the estimation from 1951 to 2003.[15] Detailed estimates and code are available on request.

[15] Reflecting the earlier starting point, my estimates for the 1950s are lower than those of Gordon (1997, p.24) and Brayton et al. (1999, p.11). However, they are higher than those of Brainard and Perry (2000, p.62) and the observations of contemporary policy-makers, as described by Christina D. Romer and David H. Romer (2002).

Chart 3: NAIRU estimated by Kalman Filter (+/- 2 RMSE)



This series has many features in common with the estimated effect of the minimum wage, shown in Chart 1. Both series are below 4 percent in the 1950s, rise above 6 percent by the late 1960s, remain high into the 1980s, fall to around 5 percent in the 1990s and finish at 4.8 percent. Variations about these trends differ considerably, but the similarity of the longer-term movements seems telling.

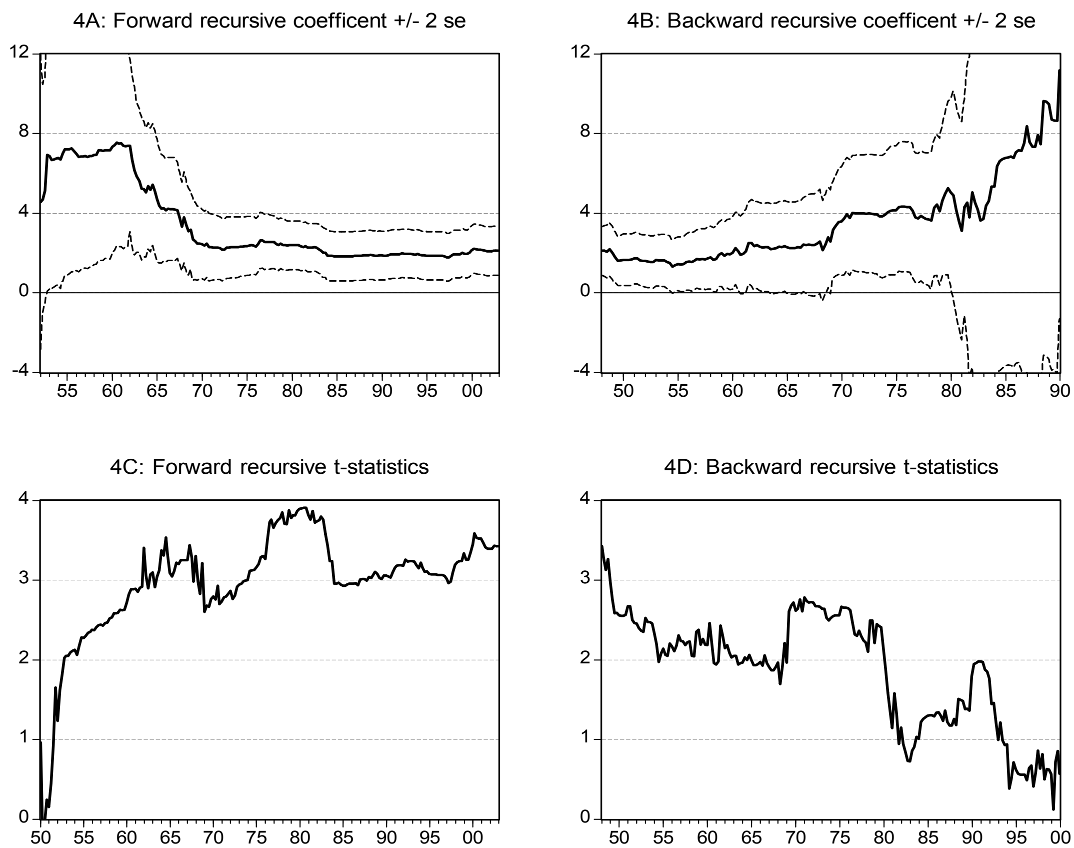
A bivariate comparison is not convincing by itself. However, it does allay concern that shocks to wages might not flow onto prices. More importantly, it shows that a leading method of estimating the NAIRU is not inconsistent with my results. Indeed, price-based estimates of the time-varying NAIRU suggest that the trends shown in Chart 1 are plausible. Not only do substantial variations in the U.S. NAIRU occur, but their magnitude and timing is roughly in line with the estimated effect of the minimum wage. Conversely, movements in the minimum wage provide an explanation of movements shown in Chart 3 – such as the 1½ percentage point decline in the NAIRU over the last two decades – which previous researchers have puzzled over and modeled as random.

6. SENSITIVITY OF WAGE EQUATION ESTIMATES

6.A. Sensitivity to sample period

Point estimates of the effect of the minimum wage appear to be relatively stable. The Andrews-Ploberger test reported in Table 1, with a *p*-value of 55 percent, indicates that differences in the coefficient across subsamples are small relative to the noise in the data. This stability is also evident in the top panels of Chart 4, which show backward and forward recursive coefficients. The estimates are positive, small (relative to wage growth) and important (relative to the effect of unemployment), regardless of when the sample starts or ends.

Chart 4: Recursive coefficients on level of minimum wage[16]



[16] Coefficients on the payroll tax and the wage freeze are constrained to equal their full-sample estimates.

Coefficient stability allays concerns about overfitting. If a variable were truly unimportant, it would be unlikely to show similar effects in two nearly independent samples. It allays concerns about omitted variable bias. Were the minimum wage proxying for an omitted variable, its coefficient would change whenever the correlation between the minimum wage and that variable changed. I investigate omitted variables in more detail in the next section. And it allays concerns about sensitivity to policy. As can be seen in Panel 4A, the effect of the minimum wage does not substantially change in the 1980s, despite introduction of a new policy of "benign neglect".

Although point estimates may not depend on the sample period, their statistical significance does. As the lower panels of Chart 4 show, the t-statistic for the minimum wage coefficient tends to increase as observations are added – though data in the middle of the sample do not seem to add much information. As Julia Campos and Neil R. Ericsson (1999) argue, rising t-statistics are further evidence against overfitting. If an unimportant variable were accidentally included, its recursive t-statistic should decline as the sample expands.

As shown in the lower right panel, 4D, backward recursive t-statistics fall below the conventional benchmark of 2 for sample periods beginning in the 1960s. This is relevant for comparisons with other research. It is common practice to estimate Phillips curves from the 1960s through to the present. So, if other researchers were to look for an effect of the minimum wage over this period, some would presumably conclude that it is statistically insignificant. Such a result is not inconsistent with the full sample results – as panel 4B shows, coefficient estimates are unaffected. However, it does point to the importance of powerful tests.

As Panel 4D also shows, when information from the 1950s and (especially) the late 1940s is taken into account, a positive effect of the minimum wage becomes increasingly clear. This period is particularly informative about the effect of the minimum wage, covering the two largest increases in the sample period, specifically the 88 percent(!) increase in 1950 and the 33 percent increase in 1956. Furthermore, as can be seen in Chart 1, observations from this period tend to lie outside the range of later periods.

Some comments seem warranted on the use of data from the 1940s and 1950s. At a practical level, taking early data into account requires that the series for prices and productivity be linked to other series before 1947. This involves effort, but the results are not very sensitive to how it is done. This is because the variation between different measures of lagged prices and trend productivity is small relative to the variation in the dependent variable and the minimum wage. For example, I extrapolate back trend productivity, a simple and conservative approach. Alternatively, using the average growth of GDP-per-worker since 1929 would increase the t-statistic on the minimum wage to 3.6.

More substantively, the early data are sometimes disregarded out of concern that they were poorly measured and contaminated by effects that are difficult to control for, such as the Korean war. However, it is not clear that these concerns matter. The structural break tests reported in Table 1 do not indicate instability. A dummy variable for the period of the Korean war is insignificant, either by itself or interacting with the minimum wage (see Table 2, below). And early residuals are not unusually large. Although several explanatory variables show unusual movements in this period, the response of wages to these movements is in line with later estimates. Other things equal, this increases confidence in the model.

Researchers analyzing current conditions tend to downweight the distant past for fear of undetected structural change. This concern is relevant if the equation is used for short-term forecasting, as wage equations frequently are. However, if the purpose is assessing a policy that is intended to apply in a wide variety of conditions, then a wide variety of data may provide a more reliable basis for any decision. A long sample seems particularly appropriate for policy analysis when the relevant policy changes have been small or infrequent.

My assessment is that the large variations in the early data are relevant and informative. Although the possibility of structural changes cannot be refuted, the data do not indicate that these have been important. Nevertheless, the weaker results obtained using conventional sample periods do justify some caution in interpreting my results.

6.B. Sensitivity to specification

An obvious reaction to my results is suspicion that the minimum wage is proxying for some omitted variable. Investigating this possibility gives some indication of the sensitivity of my estimates. It also provides guidance for future research, helps assess the adequacy of my specification and indicates the strength of evidence for other possible effects on the NAIRU – the last consideration being of wider interest.

Table 2 presents a series of F-tests on the addition of other regressors to my wage equation. Sometimes (for example, when interaction with the unemployment rate seems important) I include these contemporaneously, but mostly variables are included with lags of one and five quarters. This allows for both level and change effects, avoids reverse causation, and permits symmetric treatment of variables observed at an annual frequency. The table also shows the resulting coefficient on the level of the minimum wage and its t-statistic. My sensitivity analysis differs from that of other researchers, such as Edward E. Leamer (1983), in that I do not report sensitivity to omitting information, for example dropping or mismeasuring other regressors. Ignoring relevant evidence might well change estimates, but this is uninteresting.

Table 2: Tests of restrictions and sensitivity of minimum wage coefficient

Excluded variable	Lags	*P*-value of restriction[a]	Minimum wage coefficient	Minimum wage t-statistic	Comment (with data source in parentheses[b])
Baseline			2.1	3.4	
A. Other NAIRU determinants?					
Unemployment duration	1,5	.57	1.8	2.8	Mean unemployment duration (ES)
Unemployment benefits	1,5	.44	2.1	3.0	Pre-tax replacement rate. Unemployment insurance benefits (NIPA) divided by total unemployment (ES), all divided by compensation per hour (P&C)
Unionization rate	1,5	.20	2.4	3.5	Union members as a share of civilian employees (BLS)
Strikes	1,5	.36	1.8	2.7	Percentage of time lost due to stoppages (BLS)
Vacancy rate	0	.66	1.5	1.8	Help Wanted Index (Conference Board) divided by civilian employment (ES)
Demography	0	.51 (c)	2.1	2.7	Allowing unemployment and its demographic adjustment to enter separately (ES)
Homeownership	1,5	.64	2.2	1.9	From Residential Vacancies and Homeownership, Table 4 (Census)
Disability benefits	1,5	.38	2.5	2.6	Workers receiving disability benefits (SSA Ann. Statistical Supplement, Table 5D3), as share of labor force (ES)
Low nominal wage growth	0	.79	2.1	3.4	A dummy equal to 1 in the 23 quarters in which fitted wage growth is less than 0.75 percent a quarter
Imprisonment	1,5	.58	2.2	3.5	Share of population (Kathleen Maguire and Ann L. Pastore, 2002, Table 6.27)
Import share	1,5	.49	2.4	3.4	Ratio of imports of goods and services to gross domestic purchases (NIPA)
Immigration	1,5	.05	2.9	4.0	Immigrants admitted, less asylum seekers, divided by U.S. population (INS Yearbook Tables 1 and 4)

Excluded variable	Lags	*P*-value of restriction[a]	Minimum wage coefficient	Minimum wage t-statistic	Comment (with data source in parentheses[b])
B. Other proxies?					
Aid to families	1,5	.91	2.2	2.6	This and the next four entries represent corresponding line items from NIPA Table 2.1, all divided by gross personal income
OASDHI benefits	1,5	.40	2.7	3.4	
"Other transfer payments"	1,5	.42	2.5	3.2	
Total transfer payments	1,5	.14	2.5	3.2	
Average tax rate	1,5	.23	2.6	3.3	
C. Miscellaneous variables					
Remove inflation constraint		.26	2.3	3.4	
Remove productivity constraint		.19	2.5	3.6	
Labor share	1,5	.65	1.8	2.5	Nonfarm business sector (NIPA, P&C)
Markup	1,5	.65	2.0	3.0	Labor share, but with trend productivity instead of actual productivity
Decade dummies	0	.42	2.5	2.6	
Time and time-squared	0	.49	2.5	2.7	
Government wages	1,5	.71	2.0	2.9	NIPA index of compensation of general govt. employees divided by compensation for business employees
Unemployment above mean	0	.99	2.1	3.4	
Price wedge	1,5	.45	2.2	3.3	Inclusion of this variable is equivalent to adding product prices while preserving inflation neutrality (NIPA)
Female labor force participation	1,5	.22	1.8	2.3	Female labor force divided by non-institutional civilian population (ES)
Level of payroll tax	1,5	.40	2.6	3.3	Employer contributions to social insurance divided by compensation of employees (NIPA)
Korean War dummy	0	.45	2.2	3.5	Equal to 1 from 1950:q3 to 1953:q3
Extra lags of wages growth	3,4	.73	2.4	3.2	

Variation	Equation standard error	Minimum wage coefficient	Minimum wage t-statistic	Comment (with data source in parentheses [b])
D. Alternative measurement				
ECI instead of CPH from 1980	0.22	1.8	1.0	ECI as dependent variable; estimation from 1980:q1
ECI spliced with CPH	0.30	2.0	3.5	Splicing ECI with compensation per hour; estimation from 1948:q1
CPH instead of AHE	0.41	na	3.1	Dividing minimum wage by total compensation; coefficient is not comparable
Both the above	0.29	na	4.5	Previous two changes combined; coefficient is not comparable
CPI instead of AHE	0.41	na	2.9	Dividing minimum wage by published CPI-U
CPI instead of PCE prices	0.42	1.6	2.5	Published CPI-U, using same lag structure as baseline
Actual productivity	0.40	1.4	2.0	A 5-year average of actual productivity growth instead of kinked trend; estimation from 1953:q1

Notes:

a) The F-tests assume homoskedasticity, whereas t-statistics for the minimum wage are robust to heteroskedasticity.

b) Common data source abbreviations: (ES) Employment Situation Report, BLS; (NIPA) National Income and Product Accounts, BEA; (P&C) Productivity and Cost Report, BLS

c) *P*-value relates to null hypothesis of zero demographic effect. Relative to the "shift-share" baseline (i.e., the null hypothesis that demography has the same effect on the NAIRU as on actual unemployment), the *p*-value is 0.96.

Space limitations preclude a discussion of each line in the table, but a few general comments follow. Section A shows some variables that have been suggested as determinants of the NAIRU and for which reasonably long time series are available. These include the duration of unemployment, unemployment benefits, worker militancy (as proxied by union membership and strikes), vacancies (which should capture the effect of shifts in the Beveridge curve), demography, homeownership, disability benefits, downward nominal wage rigidity, imprisonment, import penetration and immigration. See Katz and Krueger (1999), Layard et al. (1991), Friedman (1968), the OECD (1994), David H. Autor and Mark G. Duggan (2003), Robert Shimer (1999), George A. Akerlof et al. (1996), and Andrew Oswald (1996). Contrary to suggestions in these references and elsewhere, these variables do not have clear effects on the NAIRU, judging by their insignificant *p*-values. In some cases, inclusion of new variables or the consequent shortening of the sample period reduces the statistical significance of the minimum wage. However the magnitude of the effect is little affected, changing by less than one standard error. It remains important in economic terms.

One possible interpretation of the effect of the minimum wage is that it serves as a proxy for deeper social forces such as readiness to intervene in markets or egalitarian attitudes. This hypothesis can be indirectly tested if we assume that other elements of the social safety net, such as transfer payments, also proxy for these underlying forces. Other similarly plausible proxies, presented in Section B of the table, should also have explanatory power and they should reduce the significance of the minimum wage. However, *p*-values on these other proxies are insignificant and the minimum wage coefficient is not reduced. Furthermore, these results suggest that alternative poverty reduction measures do not have the adverse unemployment effects of the minimum wage.

Section C tests the relaxation of further restrictions. Again, none suggest that the model is missing important information. Nor do they indicate that the estimated effect of the minimum wage is sensitive to specification changes.

Section D uses some alternative measures of key variables. As nested F-tests seem less interesting for these variations, I instead show the standard error of the equation. Contrary to my initial expectations, variations in measurement make more difference to estimates of the minimum wage effect, in both directions, than variations in specification or sample period. For example, the variation that shows the smallest effect – the last line in Table 2 – uses actual rather than trend productivity and begins in 1953:q1. (It retains the five-year distributed lag, which seems necessary for an adequate fit.) This reduces the coefficient on the minimum wage by a third. Even so, the effect remains statistically significant at conventional levels and the qualitative result is similar.[17]

Moreover, better measures could considerably strengthen my results. In Tulip (2000b) I splice compensation per hour with the Employment Cost Index in order to reduce high-frequency measurement error. And I divide the minimum wage by total compensation rather than by average hourly earnings, to capture differences in benefits. These refinements improve the fit of the equation and raise the t-statistic on the level of the minimum wage from 3.4 to 4.5. (The coefficient is similar once allowance is made for the different units of measurement.) However, they also raise issues of interpretation. It seems simpler to avoid these complications, though that weakens my results.

Several of the variables in the table include the lagged level of wages (the unemployment benefit replacement rate, the lagged wage share, the markup, the relative government wage, the payroll tax rate). These variables are all insignificant. Conversely, when the minimum wage is deflated by compensation per hour or consumer prices, it remains highly significant. Together these results indicate that the explanatory power of the relative minimum wage is attributable to the numerator (the minimum wage) rather than the denominator (the average wage).

The amount of sensitivity analysis that can be conducted and presented is necessarily limited and I fear that I have omitted the reader's favorite variable, for which I apologize. But based on the variations above and others I have explored, it seems that the quantitative effect of the minimum wage on nominal wage growth is only slightly sensitive to variations in model specification and the qualitative effect is robust. Of course, this conclusion is necessarily tentative and may need to be modified in light of future research.

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. Sensitivity of the unemployment effect matters for estimation, but it does not seem interesting for inference – given that my estimates are qualitatively consistent with out-of-sample evidence from other time periods, measures and countries. The chief uncertainty is whether minimum wages are an important determinant of wage growth, rather than the subsequent issue of how much unemployment is required to offset this. The focus of my sensitivity analysis reflects this.

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[17] My use of a kinked linear trend follows Akerlof et al. (1996, 2000), Gordon (1988, 1998), and Brayton et al. (1999). Although it may not alter inferences about the minimum wage, the correct treatment of productivity is an issue on which more research would be useful, given that it has large effects in many wage equations.

7. CONCLUSIONS

A substantial effect of the relative level of the minimum wage on the NAIRU helps to explain the behavior of U.S. wages, U.S. prices, the French NAIRU and international variations in unemployment. The effect is robust and clearly discernible in an equation explaining nominal wage growth in the United States. I would characterize this evidence as persuasive but not overwhelming. Reservations arise because my strongest results use an unconventionally long sample period, because strong microeconomic and theoretical underpinnings are not yet available, and because omitted variables remain a possibility. Subject to those caveats, the different sources of information seem to be consistent with a 10 percent increase in the relative minimum wage (from current U.S. levels) raising the NAIRU by about half a percentage point.

DATA APPENDIX

My data and programs are available on request and from http://www.petertulip.com. This appendix describes data sources and transformations for the wage equation. The programs provide further details including of price equations. Data include revisions up to June 2003. Many of my data adjustments are based on conversations with forecasters and others who work frequently with the data.

Wages and productivity are measured by compensation and output per hour for all persons in the non-farm business sector from the BLS Productivity and Cost report. "Trend productivity" is a linear trend fitted to the logarithm of output per hour from 1947 through 2003, with kinks at 1973:q1 and 1995:q1, which I extrapolate back to 1943. I then take a 5-year average of the quarterly change in this trend.

Consumer prices are the chain-weighted price index for personal consumption expenditures (NIPA Table 7.1). Quarterly values for annual data for 1942 to 1946 are interpolated with Eviews "quadratic match average" procedure (but then averaged again, in estimation). Prices and productivity are adjusted for the effect of unusual insurance claims in 2001:q3 and a break in 1977 (when the BLS stopped applying adjustments for recent methodological changes to the CPI).

The demographically adjusted unemployment rate is the average of unemployment rates for the five main age-sex categories, weighted by shares of the 1993 labor force, as a percentage. The series is adjusted to remove a permanent 0.08 percentage point increase in 1994 due to new survey design and a temporary 0.1 percentage point increase in 1990, phased back to 1980, arising from rebenchmarking to the Census.

The Nixon wage freeze is a dummy equal to 1 in 1971:q4 and -0.6 in 1972:q1.

The contribution of payroll taxes reflects the change in employer contributions to social insurance (NIPA Table 1.14), denoted *ec*, divided by total hours worked (from the Productivity and Cost report), denoted *hour,* as a share of projected compensation. Letting *comp* denote compensation per hour in dollars, the regressor is:

$$(ec_t/hour_t \; - \; ec_{t-1}/hour_{t-1} \times (comp_{t-2}/comp_{t-6})^{1/4}) \, / \, comp_{t-1} \;\; \times 100$$

For the minimum wage, I use whichever is the higher of the main state and Federal rate (typically that for adult men) for each state for each month and weight each state by 1996 employment shares. A time series for the Federal minimum wage and recent state rates are available at the Employment Standards Administration web site. State minima for 1981 to 1996 come from the Council of State Governments (1997) and for 1950 to 1981 from Aline O. Questor (1981,

Table 1a). 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).

Minimum wage coverage is the proportion of private sector non-supervisory employees covered by the Federal legislation, from Brown (1999, Table 1). In the absence of legislative changes, I assume coverage after 1991 remains at 86 percent.

Average hourly earnings are for all private production workers, from the BLS' Employment Situation release, which I splice with average hourly earnings in manufacturing prior to 1964. (The alternative of interpolating annual estimates poses simultaneity concerns.)

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