Chapter 1

Prices, Wages, and the U.S. NAIRU in the 1990s

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One of the most salient features of the U.S. expansion in the second half of the 1990s was the combination of low price inflation, strong real-wage growth, and low and falling unemployment. Seemingly, this runs counter to the postwar U.S. experience that periods of low unemployment and strong wage growth are associated with rising rates of inflation. This paper undertakes an empirical investigation of the extent to which changes in price-setting behavior, changes in wage-setting behavior, and fundamental changes in product and labor markets led to this happy coincidence.

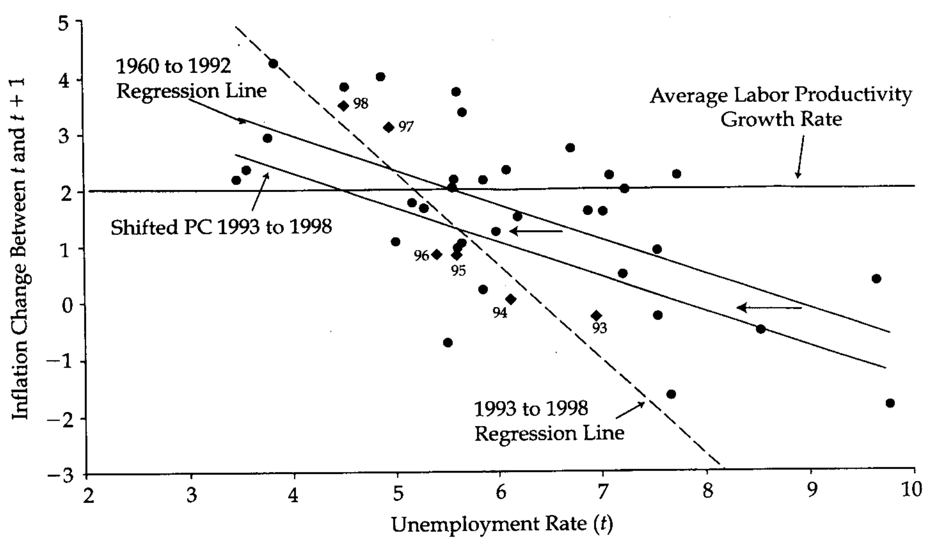
The facts are summarized in figures 1.1 and 1.2. Figure 1.1 is a scatterplot of the change in the annual rate of price inflation, as measured by the GDP deflator, against the unemployment rate in the previous year, from 1960 to 1999; for example, the point labeled *98* indicates the unemployment rate in 1998 and the change in the rate of inflation from 1998 to 1999. Figure 1.2 is a comparable scatterplot, except that the series on the vertical axis is the annual percentage growth rate of real wages, as measured by compensation per hour in the nonfarm-business sector, deflated by the GDP deflator. A regression line estimated using data from 1960 to 1992 is plotted in both figures. These regression lines are simple estimates of the price and wage Phillips curves. The NAIRU (nonaccelerating inflation rate of unemployment) is defined to be the value of the unemployment rate at which the price regression predicts no change in inflation, which corresponds to the intersection of the regression line and the horizontal line in figure 1.1. Alternatively, the NAWRU (nonaccelerating wage rate of unemployment), the wage-based NAIRU, is the value of the unemployment rate at which the wage-regression line predicts real-wage growth that coincides with the growth in labor productivity, which is given by the intersection of the regression line and the horizontal line in figure 1.2.

Three features are evident from these scatterplots. First, both the wage and the price Phillips curves reflect a negative correlation between the unemployment rate in one year and inflation in the next: the correlation for the 1960 to 1992 sample is -0.55 in figure 1.1 and -0.65 in figure 1.2. Second, the data for 1993 to 1999 (highlighted in the figures) are peculiar, relative to the earlier data:

FIGURE 1.1 / Price Inflation and Unemployment

Sources: Bureau of Labor Statistics and authors' calculations.
Note: PC = Phillips curve.

FIGURE 1.2 / Real-Wage Inflation and Unemployment



Sources: Bureau of Labor Statistics, Bureau of Economic Analysis, and authors' calculations.
Note: PC = Phillips curve.

although unemployment fell from 7.5 percent in 1992 to 4.1 percent in 1999, the rate of price inflation was essentially constant over this period (it fell by an average of 0.1 percent per year). Third, from 1993 to 1999, real wages increased substantially: real wages grew by an average of 1.5 percent over this period, consistent with little, if any, shift in the NAIRU in this wage scatterplot and, if anything, a steeper regression line for the wage Phillips curve.

The theories that have been proposed to explain these events fall into two groups: those proclaiming that "the Phillips curve is alive and well but . . ." and those proclaiming that "the Phillips curve is dead." Most of the proposed theories are in the "alive and well but . . ." group. According to these theories, the price Phillips curve–the regression line in figure 1.1–continues to have a negative slope but has been shifting inward. Such a shift is indicated by the arrow and new line in figure 1.1. Similarly, the strong growth of real wages in the late 1990s in figure 1.2 is attributed to the surge in productivity: workers are reaping the rewards of using more powerful tools. The differences among these theories arise in the particulars of their explanations of the inward shift of the price-inflation Phillips curve: some focus on the price-setting behavior of firms, others focus on labor markets, while still others suggest that we have simply been the lucky recipients of favorable supply shocks (falling energy prices and favorable terms-of-trade shocks).

The theories that focus on pricing behavior have several variants. One is that globalization has increased competition in the product market, thereby squeezing markups and yielding one-time reductions in markups and prices (for example, Brayton, Roberts, and Williams 1999). Similar arguments can be made about the possible effect of the Internet on price competition for some goods. A different argument is that the credibility of the commitment of the Federal Reserve Board to controlling inflation has increased and that this has had the effect of reducing expected inflation, which in turn moderates actual price increases posted by producers.

The theories that focus on labor markets suggest that the source of the inward shift in the Phillips curve lies in a decline in the natural rate of unemployment. Several such theories are surveyed and analyzed empirically in Lawrence Katz and Alan Krueger (1999). Some emphasize changes in how people look for work (using temporary-help firms, the Internet, etc.). Others emphasize changes in the composition of the workforce, including the aging of the workforce as the baby boom enters an age traditionally associated with high degrees of labor-force attachment, the entry of "welfare mothers" into the workforce as a consequence of welfare reform, and the removal of many marginal workers from the workforce either because of incarceration (Katz and Krueger 1999) or because of relaxed social security disability insurance provisions (Autor and Duggan 2000).

Finally, some of these theories stress the role of good luck. For much of the 1990s, energy prices were declining, and the United States enjoyed a strong dollar. Robert Gordon (1998) explored these sources in detail, concluding that they explain part, but far from all, of the price-inflation–unemployment puzzle of the 1990s.

In contrast, "the Phillips curve is dead" theories interpret the 1990s, not as an inward drift in the Phillips curve, but rather as a fundamental change in the relation between unemployment and inflation. According to these theories, it is the slope of the Phillips curve that has shifted, not the intercept: the Phillips curve now is the dotted line in figure 1.1, which was fit to the data from 1993 to 1999. This curve has a slope of 0. This more radical interpretation requires more radical theories.

The popular-press versions of these theories stress that increased price competition in the new economy prevents firms from responding to market tightness by increasing prices, thereby eliminating any relation between measures of aggregate activity, such as unemployment, and changes in the rate of price inflation. Subtler versions of these theories involve nonlinearities in firm behavior when inflation is low. George Akerlof, William Dickens, and George Perry (1996) suggest that reluctance by firms to give negative nominal-wage cuts means that steady-state hiring depends on the rate of inflation; in particular, the equilibrium unemployment rate falls when inflation falls. John Taylor (2000) develops a different theory of the state dependence of the NAIRU; in his model, low inflation itself leads firms to expect reduced pricing power, which in turn contributes to reduced inflation and reduces the sensitivity of inflation to growth in demand. Akerlof, Dickens, and Perry (2000) provide a model of price setting in which some firms find it convenient to predict no inflation as long as inflation is low, permitting the unemployment rate to be persistently low without igniting inflation. Empirically, this is the same thing as the NAIRU falling when the inflation rate gets low. In all three of these models, the NAIRU is not *permanently* low; rather, its low value is contingent on the monetary authority holding down inflation.

This paper has two objectives. The first, more modest one is to document the shifts in figures 1.1 and 1.2. To a considerable extent, this entails updating earlier estimates of Phillips curves and NAIRUs along the lines of Staiger, Stock, and Watson (1997a, 1997b) and Gordon (1998). The second, more ambitious objective is to provide new evidence, based on quarterly macro data and on a panel of annual data for U.S. states from 1979 to 1999, that will help us parse the theories outlined in this introductory section or, at least, rule out some families of theories.

We address these two objectives by asking three specific questions. First, did the Phillips curve break down in the 1990s, or did it simply shift with a new and evolving NAIRU? That is, which class of theories—"the Phillips curve is dead" or "the Phillips curve is alive and well but . . ."—has more empirical support? We conclude that the weight of the evidence suggests that the price Phillips curve has shifted in, not flattened out, supporting the "alive and well but . . ." group.

This leads to the second question: Why has the price Phillips curve shifted in? That is, does the empirical evidence help distinguish between the many theories of the inward drift in the price Phillips curve? In our view, the weight of the empirical evidence points toward explanations that involve special features in labor markets. The macro evidence suggests that changes in price-setting behavior cannot explain the broad stability of the relation between price inflation and measures of economic activity. Rather, the explanation for the shifting unemployment Phillips curve seems to lie in declines in the univariate trend rate of unemployment.

The third question, then, is whether labor-productivity gains during the 1990s can explain the apparently aberrant recent behavior of real wages in figure 1.2. That is, is the wage Phillips curve as resilient empirically as the price Phillips curve once we have accounted for productivity? Our answer is yes: adjusting for trend labor-productivity gains accounts for the discrepancies that otherwise appear between the price and the wage Phillips curves.

In short, once one allows for the *univariate* trends in the unemployment rate and the rate of productivity growth, the 1990s present no wage or price puzzles. Backward-looking price Phillips curves are stable when the unemployment rate is specified as a gap, that is, as a deviation from its univariate trend value. Similarly, wage Phillips curves are stable when wages are adjusted for changes in trend productivity growth and when the regressions are specified using activity gaps. This implies that theories of the 1990s that focus on favorable supply shocks, changes in the pricing power of firms and markups, or changes in the negotiating power of labor all miss the mark, for they imply persistent errors and/or coefficient instability that we fail to find. Rather, the evidence points to underlying economic forces that change the univariate trends of the unemployment rate and the growth rate of productivity. Unfortunately, our regressions using the state data fail to isolate any economic or demographic determinants of the trend unemployment rate.

The plan of the paper is as follows. The first four sections analyze quarterly U.S. macro data from 1960 to 2000. We begin in the first section by estimating the long-run trends in the macro data and discussing how we estimate output gaps. The second section addresses issues of econometric specification and estimation of the price and wage Phillips curves and associated time-varying NAIRUs (TV-NAIRUs). These Phillips curves are specified using output gaps, which are the difference between the output measure and its low-frequency univariate trend component. As Robert Hall (1999) and T. Cogley and T. Sargent (forthcoming) argue, the low-frequency trend component can be thought of as an estimate of the natural rate of unemployment; thus, this approach allows separate identification of the NAIRU and the natural rate. The third section reports empirical price Phillips relations estimated both with the unemployment-rate gap and with gaps based on other measures of economic activity. Consistent with the findings reported in Douglas Staiger, James Stock, and Mark Watson (1997b), Stock (1998), and Stock and Watson (1999b), we find stability and predictive content in these broader measures, which suggests that "the Phillips curve is dead" theories are premature. In the fourth section, we turn to wage Phillips curves and examine the role of productivity gains in explaining the recent rise in real wages.

The next four sections focus on the state panel data. The use of state-level data has been limited (for notable exceptions, see Katz and Krueger (1999) and Lerman and Schmidt 1999), and we are able to consider a large number of new variables and, accordingly, use the state data to examine the various theories. Specifications and econometric issues, including our instrumental-variables (IV) method for alleviating errors-in-variables bias arising from using the state data, are discussed in the fifth section. The data set is described in the sixth section, and benchmark results are presented in the seventh section. The eighth section reports the results of using additional variables to explore the stability of the Phillips curve and to examine theories about sources of shifts in the NAIRU. Conclusions are summarized in the final section.

A remark on terminology is in order before proceeding. In conventional usage, the *NAIRU* is the rate of unemployment consistent with price inflation remaining constant; the *NAWRU* is the rate of unemployment consistent with wage inflation remaining constant; and the *NAIRCU* is the rate of capacity utilization consistent with price inflation remaining constant. In this paper, we consider both wage and price inflation as well as other activity indexes, including building permits and demographically adjusted unemployment. We could, then, report TV-NAWRCUs, TV-NAIRBPs, TV-NAWRDUs, and so on. But we do not find these acronyms helpful. Instead, we shall call them all *TV-NAIRUs* and, when needed, add specificity through the use of adjectives.

TRENDS IN THE MACRO DATA

Method for Estimating Univariate Trends and Constructing Gaps

Let y_t be a quarterly time series, and let y_t^* denote its trend. Unless explicitly noted otherwise, y_t^* is estimated by passing y_t through a two-sided low-pass filter, with a cutoff frequency corresponding to fifteen years. Essentially, this estimates y_t^* as a long two-sided weighted moving average of y_t with weights that sum to 1. Estimates of the trend at the beginning and the end of the sample are obtained by extending (padding) the series with autoregressive forecasts and backcasts of y_t, constructed from an estimated AR(4) model (with a constant term) for the first-difference of y_t. The "gap" value of y_t, y_t^g, is defined to be the deviation of y_t from its trend value; that is, $y_t^g = y_t - y_t^*$. Thus, the trend value of the unemployment rate is the value of the unemployment rate resulting from the low-pass filter, and the unemployment-rate gap is the difference between the actual unemployment rate and the long-run trend in unemployment.

Description of the Aggregate U.S. Data

The U.S. data are quarterly from 1959:1 to 2000:2. The primary price measure is the GDP deflator, but, in our sensitivity analysis, we also consider the personal-consumption-expenditure (PCE) deflator, the CPI, and the deflator for the nonfarm-business component of GDP. All rates of inflation are computed as $\pi_t = 400 \ln(P_t/P_{t-1})$, where P_t is the level of the price index in quarter t.

Several measures of wages are used. Our primary measure is compensation per hour in the nonfarm-business sector. For sensitivity checks, we also consider the employment cost index (ECI)—both total compensation and wages and salaries only—average hourly earnings of nonagricultural production workers, and compensation per hour in manufacturing. Wage-growth rates are computed as $\omega_t = 400 \ln(W_t/W_{t-1})$, where W_t is the level of the wage index.

Labor productivity is measured by output per hour of all workers in the nonfarm-business sector, except when we consider compensation per hour in manufacturing, in which case labor productivity is measured by output per hour of all workers in manufacturing.

Economic activity is variously measured by the total unemployment rate, a demographically adjusted unemployment rate, the rate of capacity utilization, and housing starts (building permits). The demographically adjusted unemployment rate was constructed as a weighted average of the unemployment rates for fourteen age-gender categories (ages sixteen to nineteen, twenty to twenty-four, twenty-five to thirty-four, thirty-five to forty-four, forty-five to fifty-four, fifty-five to sixty-four, and sixty-five and over, each by gender), weighted by the shares of each age group in the 1985 labor force.

Supply-shock variables in the Phillips curve regressions are Gordon's (1982) price-control series, the relative price of food and energy, and exchange rates.

Data sources for all series are given in the data appendix.

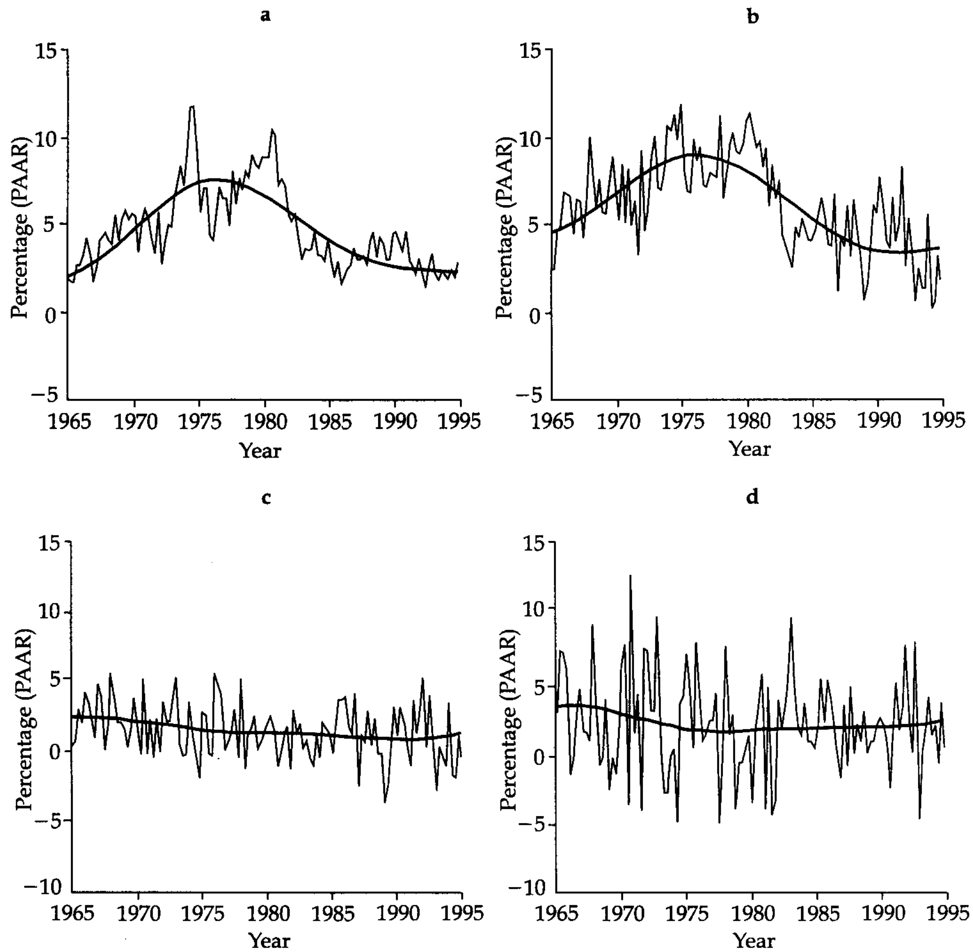
Low-Frequency Properties of the Data

Figure 1.3 presents quarterly time-series data and their estimated trends for (*a*) price inflation (the GDP deflator), (*b*) wage inflation (compensation per hour), (*c*) real-wage growth (compensation per hour deflated by GDP-deflator inflation), (*d*) labor-productivity growth, (*e*) the unemployment rate, (*f*) the demographically adjusted unemployment rate, (*g*) building permits (housing starts), and (*h*) the rate of capacity utilization. The "gaps" of each of these variables are the difference between the quarterly data and their estimated trends. Table 1.1 shows the sample mean of each of these series over each of the four decades in the sample and, as a measure of persistence in the series, a 95 percent confidence interval for the largest root in a univariate autoregression with six lags.

Figure 1.3 and table 1.1 show several important features of these data. First, consider wages, prices, and productivity. Figures 1.3*a* and 1.3*b* show substantial low-frequency (trend) variability in price and nominal-wage inflation. As shown in table 1.1, this low-frequency variability leads to confidence intervals for the largest AR root that range from 0.90 to 1.02 (intervals that, notably, include a unit AR root). Nominal-wage growth less productivity growth is also persistent: the confidence interval for its largest AR root is 0.81 to 1.00.

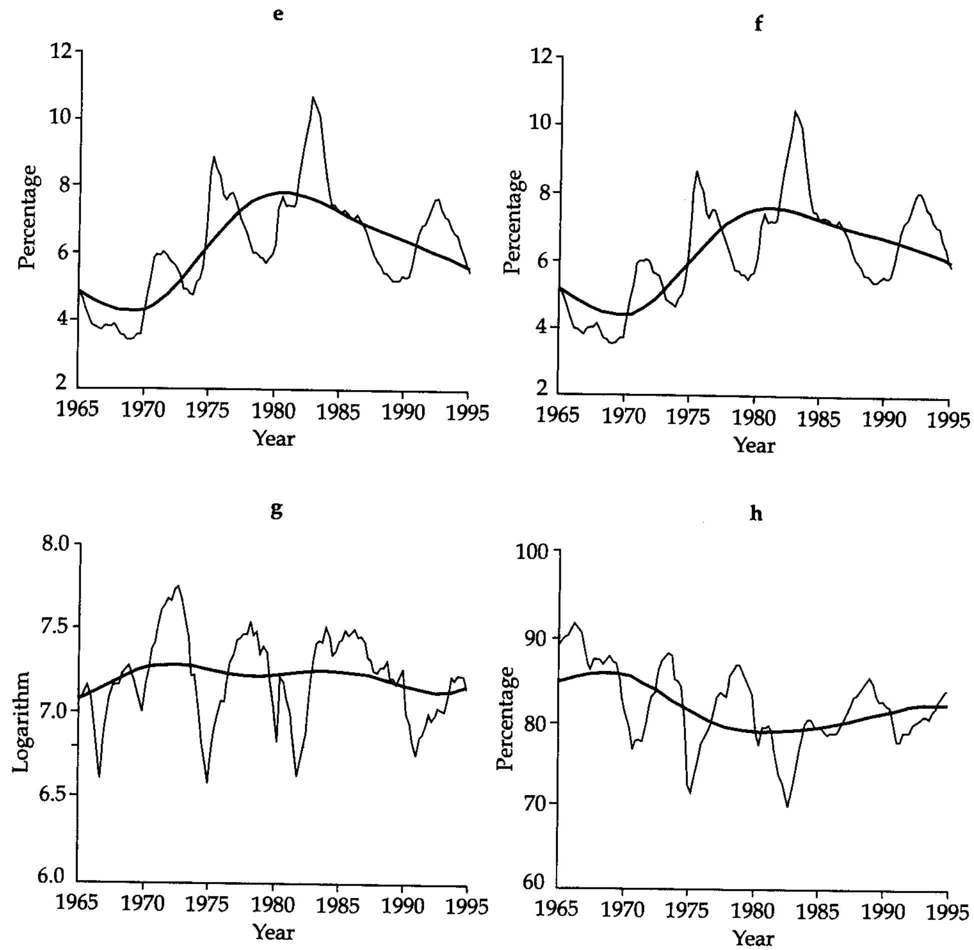
In contrast, real-wage growth and, especially, real-wage growth less produc-

FIGURE 1.3 / Macro Series and Their Trend Values



tivity growth are considerably less persistent, and 95 percent confidence intervals for their largest AR roots do not include unity. Still, the decade-long averages in real wages show considerable variability, and the ratio of the largest to the smallest decadal average varies by a factor of more than 2.5. Table 1.1 shows that these decade-long changes in real-wage-growth rates are broadly consistent with movements in the growth of labor productivity: real-wage growth and labor-productivity growth were both high in the 1960s, low in the 1980s, and so on. The relation is stronger when wages, prices, and productivity all pertain to the nonfarm-business sector than when the GDP deflator is used to construct

FIGURE 1.3 / *(Continued)*



Sources: Bureau of Labor Statistics, Bureau of Economic Analysis, and authors' calculations.
Notes: a, Price inflation; *b,* Wage inflation; *c,* Real-wage inflation; *d,* Productivity growth; *e,* The unemployment rate; *f,* The demographically adjusted unemployment rate; *g,* Building permits; *h,* Capacity utilization.

real wages. Average real-wage growth adjusted for productivity changes little over the decades in the sample.

Formally, the hypothesis of a unit root in productivity is rejected, which, taken literally, indicates that productivity growth is stationary. This characterization, however, does not allow for the possibility of slowly changing mean productivity-growth rates that lie at the heart of the new-economy debate. From a sta-

TABLE 1.1 / Descriptive Statistics for Trend Characteristics of the Data

	Sample Mean				95 Percent CI
Series	60:1–69:4	70:1–79:4	80:1–89:4	90:1–99:4	for the Largest AR Root
Price Inflation	2.47	6.50	4.46	2.23	.90 to 1.02
Nominal Wage Growth Rate	4.95	8.05	5.36	3.72	.91 to 1.02
Nominal Wage-Productivity Growth Rate	2.22	6.09	3.94	1.70	.81 to 1.00
Real Wage[1] Growth Rate	2.48	1.55	.90	1.49	<.93
Real Wage[2] Growth Rate	2.86	1.81	1.15	1.71	<.89
Real Wage[1]-Productivity Growth Rate	−.25	−.40	−.52	−.52	<.77
Real Wage[2]-Productivity Growth Rate	.14	−.15	−.27	−.30	<.77
Productivity Growth Rate	2.72	1.96	1.42	2.02	<.80
Real Wage[1]/Productivity (log)	−.029	−.067	−.080	−.138	.97 to 1.03
Real Wage[2]/Productivity (log)	−.014	−.010	−.012	−.052	.91 to 1.02
Unemployment Rate	4.78	6.21	7.27	5.76	.89 to 1.02
Demographically adjusted unemployment rate	5.04	6.06	7.26	6.11	.86 to 1.01
Building Permits (log)	7.07	7.31	7.24	7.17	<.89
Capacity-Utilization Rate	85.0	81.6	79.0	80.9	<.93

Note: Columns 1 to 4 show the sample means of the series listed in the stub column over the sample period indicated. The final column shows the 95 confidence interval (CI) for the largest root in a univariate AR(6) model (including the constant). The sample period for the regression was 1960:1 to 2000:2. The confidence interval was computed using the approximation developed in Stock (1991), for highly persistent series. Several of the series were not very persistent, and Stock's method could be used to compute only an upper confidence bound. Real Wage[1] uses price inflation computed from the GDP deflator, and Real Wage[2] uses price inflation computed from the price deflator for the nonfarm business sector.

tistical point of view, if there is a highly persistent component of productivity growth but its variance is small relative to variations induced by cyclical movements and measurement error, then it will be difficult to detect, and the series can spuriously appear to be stationary.

These results are consistent with price inflation and nominal-wage inflation adjusted for productivity growth sharing a common stochastic trend that disappears from real-wage growth adjusted for productivity growth. In the terminology of integration and cointegration, this suggests that price inflation is I(1) (that is, is integrated of order 1), wage inflation less productivity growth is I(1), and real-wage growth less productivity growth is I(0); that is, price inflation is cointegrated with wage inflation less productivity growth, with a cointegrating vector of $(1, -1)$.[1]

Consistent with this specification, the *level* of productivity-adjusted real wages (equivalently, the "markup" or "labor's share") appears to be I(1). When real wages are computed using the GDP deflator, there is a marked downward trend in the markup, but, when the GDP deflator is replaced by the nonfarm-business deflator, much of the trend disappears. In either case, the series is very persistent, and a unit autoregressive root cannot be rejected.

The unemployment rate, building permits (housing starts), and capacity utilization are shown in figures 1.3*e* to 1.3*h*. The unemployment-rate trend exhibits great variability (figure 1.3*e*), most of which remains in the demographically adjusted unemployment rate (figure 1.3*f*). In contrast, the trends in building permits and capacity utilization (figs. 1.3*g* and 1.3*h*) show much less variability. Table 1.1 indicates that the unemployment-rate series are much more persistent than the building-permits and capacity-utilization series: unit autoregressive roots cannot be rejected for either unemployment series, but they can be rejected for both building permits and capacity utilization.

In summary, these statistics suggest that price inflation and nominal-wage inflation, adjusted for productivity growth, are cointegrated. Real wages and productivity growth move together at low frequencies, although these movements are small in magnitude compared with the noise and cyclical movements in these series. The unemployment rate and the demographically adjusted unemployment rate appear to be I(1), but capacity utilization and building permits are I(0).

SPECIFICATION AND ESTIMATION OF MACRO PRICE AND WAGE EQUATIONS

Specification

Our specifications and estimation methodology follow along the lines of Robert Gordon (1982), Robert King, James Stock, and Mark Watson (1995), Douglas Staiger, James Stock, and Mark Watson (1997a), and Robert Gordon (1998), with some modifications.

Because prices and wages are codetermined, and because we will examine both price and wage Phillips curves, it is useful to consider these curves as a system. Our discussion of trends in the last section suggests that it is fruitful to treat wage inflation and price inflation as a cointegrated system, with each variable being integrated of order 1 and having the single cointegrating vector implying that real-wage growth net of productivity growth is integrated of order 0.

MOTIVATION FROM A SYSTEM WITHOUT LAG DYNAMICS Let π_t denote the rate of price inflation, let ω_t be the rate of nominal-wage inflation (the growth rate of nominal wages), and let θ_t be the growth rate of labor productivity (all expressed in units of percentage annual growth rates). Let x_t be a demand-gap variable, for example, the output gap or the unemployment gap constructed using the method outlined earlier. Let Z_t be a vector of mean 0 variables representing observable supply shocks (such as shifts in the relative prices of food and energy) that might affect wage and price setting and thus might enter either the wage or the price equations.

The price equation relates the deviation of future inflation from its expectation to the activity gap and supply shocks. Ignoring lags for the moment, this is

$$\pi_{t+1} - \pi^e_{t+1} = \mu_\pi + \beta_\pi x_t + \gamma_\pi Z_t + \nu_{\pi t+1}, \tag{1.1}$$

where π^e_{t+1} is the inflation in period $t + 1$ that is expected as of period t, μ_π, β_π, and γ_π are unknown coefficients, and $v_{\pi t+1}$ is an error term.

Implementation of equation 1.1 requires specifying inflationary expectations. Following an old convention (see Gordon 1990, 1998; and Fuhrer 1995), we restrict attention to the random-walk model of expectations, with the result that $\pi^e_{t+1} = \pi_t$ and $\pi_{t+1} - \pi^e_{t+1} = \Delta\pi_{t+1}$, where $\Delta\pi_{t+1} \equiv \pi_{t+1} - \pi_t$. Making this modification, we have

$$\Delta\pi_{t+1} = \mu_\pi + \beta_\pi x_t + \gamma_\pi Z_t + \nu_{\pi t+1}. \tag{1.2}$$

The wage equation is obtained similarly. Again ignoring lags, we have

$$\omega_{t+1} - \omega^e_{t+1} = \mu_\omega + \beta_\omega x_t + \gamma_\omega Z_t + \nu_{\omega t+1}. \tag{1.3}$$

Various approaches are available for modeling expected nominal wages. We model expected wage inflation as the sum of expected price inflation and expected productivity growth, that is, $\omega^e_{t+1} = \pi^e_{t+1} + \theta^e_{t+1}$. As in the price equation, we suppose that $\pi^e_{t+1} = \pi_t$. If productivity growth is a random walk, then we can let $\theta^e_{t+1} = \theta_t$. However, productivity growth has a cyclical component, so an alternative method used by Gordon (1998) is to model $\theta^e_{t+1} = \theta^*_t$, where θ^*_t is trend productivity growth. We will use this latter approach as the base specification, but we will also report results that are based on the alternative specification in which $\theta^e_{t+1} = \theta_t$. This leads to a specification of the wage equation of

$$\omega_{t+1} - \theta^*_t - \pi_t = \mu_\omega + \beta_\omega x_t + \gamma_\omega Z_t + \nu_{\omega t+1}. \tag{1.4}$$

INCORPORATION TO LAG DYNAMICS The specifications 1.2 and 1.4 omit lag dynamics. Our treatment of dynamics is motivated by the observation, made earlier, that nominal-wage growth less productivity growth, or less trend productivity growth, appears to be cointegrated with price inflation. That is, $\omega_{t+1} - \theta^*_t$ and π_t are arguably cointegrated. This leads to the triangular representation of cointegrated variables in which x_t and Z_t are treated as exogenous variables:

$$\begin{aligned}\Delta\pi_{t+1} = \mu_\pi &+ \alpha_{\pi\pi}(L)\Delta\pi_t + \alpha_{\pi\omega}(L)(\omega_t - \theta^*_{t-1} - \pi_{t-1}) + \beta_\pi x_t \\ &+ \alpha_{\pi x}(L)\Delta x_t + \gamma_\pi Z_t + \nu_{\pi t+1},\end{aligned} \tag{1.5}$$

$$\omega_{t+1} - \theta_t^* - \pi_t = \mu_\omega + \alpha_{\omega\pi}(L)\Delta\pi_t + \alpha_{\omega\omega}(L)(\omega_t - \theta_{t-1}^* - \pi_{t-1}) + \beta_\omega x_t + \alpha_{\omega x}(L)\Delta x_t + \gamma_\omega Z_t + \nu_{\omega t+1}, \tag{1.6}$$

where μ_π, β_π, and so on are coefficients, and $\alpha_{\pi\pi}(L)$ and so on are lag polynomials. Specifications 1.5 and 1.6 allow lagged effects of x_t but, following the literature, not of the supply-shock variable Z_t.

These two specifications form the basis for our time-series analysis. The price equation differs from most Phillips curve specifications because it includes a term allowing feedback from real wages net of productivity to future price changes. The wage equation also allows for feedback from price changes to future wage changes. Our motivation for these equations has been to move from the static system equations 1.2 and 1.4 using the tools of cointegration theory. Note, however, that our resulting equations are the same as the general specification considered by Gordon (1998, equations 7 and 8).[2]

An alternative specification that we explore in the empirical section adds a lag of the *level* of productivity-adjusted real wages ($\ln[W_{t-1}/P_{t-1}] - \ln[\text{productivity}_{t-1}]$) to the right-hand side of equations 1.5 and 1.6. This specification, a version of which goes back to the classic paper by Sargan (1964), is appropriate when this term is I(0). As the analysis in the last section suggested, this assumption seems at odds with the data used here, but versions of the specification have been used for both wage and price equations using data from the United States and other countries (see Barlow and Stadler, n.d.; Blanchflower and Oswald 1994; Brayton, Roberts, and Williams 1999; and Holden and Nymoen 1999). Olivier Blanchard and Lawrence Katz (1997) (and the references cited therein) contains a useful discussion of this specification as it applies to the wage Phillips curve, and we will discuss this issue more in the context of the state Phillips curves specified in our fifth section.

Estimation of TV-NAIRUs

Specifications 1.5 and 1.6 use the "gap" variable x_t, constructed as the difference between the activity variable and its univariate trend, but what should appear in the Phillips curve is the deviation between the variable and the NAIRU. So, if the univariate trend and the variable's NAIRU are different, then 1.5 and 1.6 should include another term that captures this difference. We model the difference between the NAIRU and the univariate trend as a time-varying intercept in these Phillips curves and estimate this difference from estimates of the time-varying intercept.

To make this clear, consider the system in which the activity measure is the rate of unemployment, u_t, and let u_t^N denote the possibly time-varying NAIRU. If the NAIRU does not equal the univariate trend u_t^*, then equation 1.5 is properly specified as

$$\begin{aligned}\Delta\pi_{t+1} &= \mu_\pi + \alpha_{\pi\pi}(L)\Delta\pi_t + \alpha_{\pi\omega}(L)(\omega_t - \theta^*_{t-1} - \pi_{t-1}) \\ &\quad + \beta_\pi(u_t - u^N_t) + \alpha_{\pi u}(L)\Delta(u_t - u^N_t) + \gamma_\pi Z_t + \nu_{\pi t+1} \\ &\cong (\mu_\pi + \beta_\pi[u^*_t - u^N_t]) + \alpha_{\pi\pi}(L)\Delta\pi_t \\ &\quad + \alpha_{\pi\omega}(L)(\omega_t - \theta^*_{t-1} - \pi_{t-1}) + \beta_\pi u^g_t + \alpha_{\pi u}(L)\Delta u_t \\ &\quad + \gamma_\pi Z_t + \nu_{\pi t+1},\end{aligned} \tag{1.7}$$

where $u^g_t = u_t - u^*_t$ is the unemployment gap, and the second equation makes the approximation that, because U^N_t is slowly varying, the term $\alpha_{\pi u}(L)\Delta u^N_t$ is negligible. Thus, to the extent that the univariate trend in unemployment u^*_t differs from the NAIRU u^N_t, the gap specification 1.5 will have a time-varying intercept. An identical argument applies to the wage equation 1.6.

This reasoning leads to a modification of the system equations 1.5 and 1.6, in which the intercepts are allowed to vary over time:

$$\begin{aligned}\Delta\pi_{t+1} &= \mu_{\pi t} + \alpha_{\pi\pi}(L)\Delta\pi_t + \alpha_{\pi\omega}(L)(\omega_t - \theta^*_{t-1} - \pi_{t-1}) + \beta_\pi u^g_t \\ &\quad + \alpha_{\pi x}(L)\Delta x_t + \gamma_\pi Z_t + \nu_{\pi t+1},\end{aligned} \tag{1.8}$$

$$\begin{aligned}\omega_{t+1} - \theta^*_t - \pi_t &= \mu_{\omega t} + \alpha_{\omega\pi}(L)\Delta\pi_t + \alpha_{\omega\omega}(L)(\omega_t - \theta^*_{t-1} - \pi_{t-1}) \\ &\quad + \beta_\omega u^g_t + \alpha_{\omega x}(L)\Delta x_t + \gamma_\omega Z_t + \nu_{\omega t+1}.\end{aligned} \tag{1.9}$$

If the slope coefficients are stable, any intercept drift in these equations arises from a departure of the NAIRU from the trend unemployment rate.

Our method for estimating the intercept drift follows King, Stock, and Watson (1995), Staiger, Stock, and Watson (1997a), and Gordon (1997, 1998) and adopts an unobserved-components model for the intercept, in which the intercept follows a random walk:

$$\mu_{\pi t+1} = \mu_{\pi t} + \eta_{\pi t+1}, \quad \text{where } \eta_{\pi t+1} \text{ is i.i.d. } N(0, \sigma^2_{\eta_\pi}); \tag{1.10}$$

$$\mu_{\omega t+1} = \mu_{\omega t} + \eta_{\omega t+1}, \quad \text{where } \eta_{\omega t+1} \text{ is i.i.d. } N(0, \sigma^2_{\eta_\omega}). \tag{1.11}$$

The random-walk specification is a flexible way to track smooth changes in the intercept. The initial condition for the random walk is identified by the unconditional means of the regressors, so we construct the regressors to have mean 0 and initialize the random walk at 0.

According to the system 1.8 and 1.9, time variation in the wage- and price-equation intercepts arises from changes in $u^*_t - u^N_t$, and this means that the innovations $\eta_{\pi t+1}$ and $\eta_{\omega t+1}$ should be the same. We shall examine this by estimating the intercept drift separately for the price and the wage equations and comparing the results. In addition, we shall (separately) test the hypotheses that $\sigma^2_{\eta_\pi} = 0$ and $\sigma^2_{\eta_\omega} = 0$ using the QLR or sup-Wald test (Quandt 1960; Andrews 1993). The parameters $\sigma^2_{\eta_\pi}$ can be estimated by maximum likelihood, but the maximum likelihood estimator has a distribution that piles up at 0 when these are small and is thus unsatisfactory. Instead, we construct confidence intervals and median-unbiased estimates of $\sigma^2_{\eta_\pi}$ and $\sigma^2_{\eta_\omega}$ using the methods in Stock and

Watson (1998), as discussed in Staiger, Stock, and Watson (1997a) and Stock (1998, in press).

The estimate of the NAIRU $\hat{u}_t^N$ based on one of these estimated equations is obtained by combining the univariate drift and the intercept drift. Because $\mu_{\pi t} = \beta_\pi(u_t^* - u_t^N)$ (with mean 0 regressors), we have the estimator,

$$u_{t|T}^N = u_t^* - \frac{\eta_{\pi t|T}}{\hat{\beta}_\pi}, \qquad (1.12)$$

where $\hat{\beta}$ is the estimator of β_π, and $\mu_{\pi t}|_T$ is the estimator of $\mu_{\pi t}$ obtained from the Kalman smoother implemented with the estimated parameters of the system.

We have motivated this treatment of the parameter drift by observing that we want a consistent framework that is flexible enough to handle activity measures with quite different trends. However, this formulation has two additional advantages. First, it allows separate identification of the univariate trend and the NAIRU. Second, since much of the time variation in the NAIRU is likely to be associated with changes in the trend unemployment rate, the method can be viewed as a device akin to prewhitening to obtain more precise estimates of the TV-NAIRU.

MACRO ESTIMATES OF PRICE PHILLIPS CURVES

Benchmark Price Regressions

Benchmark estimates of regressions of the form 1.8, using various activity gaps, are reported in table 1.2. The specifications include standard supply-shock variables (Gordon's [1982] series for wage and price controls and the relative price of food and energy). For comparability to conventional specifications, these specifications do not include the error-correction term (real wages less trend productivity, $\omega_t - \theta_{t-1}^* - \pi_{t-1}$) or its lags as regressors; these are included in results reported in the next section. The first row reports the estimated value of the coefficient on the level of the activity gap (which is the sum of the coefficients in a specification that uses the current and lagged gaps), its standard error, and the *p*-value for the sup-Wald statistic testing the stability of this coefficient. The second block of entries reports the trend value of the activity measure, and the third block of entries reports the estimated TV-NAIRU (the sum of the univariate trend and the estimated deviation of the trend arising from intercept drift). Standard errors for the NAIRU, and for its change since 1992, are computed using the Kalman smoother standard-error formula and do not incorporate estimation error, which would increase them. The final row reports the median-unbiased estimate of the standard deviation of the change in the intercept; if the population counterpart of this coefficient is 0, there is no parameter drift, so the TV-NAIRU for that activity measure equals its univariate trend.

Four results are notable. First, the slope coefficient shows the procyclical nature of the change in inflation and is statistically significant in each of these

TABLE 1.2 / Phillips Curve Estimates from Macroeconomic Data: Price Inflation Equation
$\Delta\pi_{t+1} = \beta_\pi(u_t - u_t^N) + \alpha_{\pi\pi}(L)\Delta\pi_t + \alpha_{\pi x}(L)\Delta u_t + \gamma_\pi Z_t + \nu_{\pi t+1}$

	Civilian Unemployment Rate	Demographically Adjusted Unemployment Rate	Capacity Utilization	Building Permits
Phillips curve slope (SE) [stability *p*-value]	−.28	−.26	.09	2.18
	(.10)	(.09)	(.03)	(.44)
	[.06]	[.13]	[.02]	[.16]
Trend values				
1970	4.33	4.39	85.51	7.27
1980	7.79	7.55	78.74	7.23
1990	6.40	6.66	80.84	7.18
2000	4.48	4.83	81.02	7.31
Change 1992 to 2000	−1.60	−1.59	−.57	.17
NAI trend values (SE)				
1970	4.62	4.66	84.86	7.23
	(.41)	(.44)	(1.25)	(.05)
1980	7.73	7.50	79.11	7.25
	(.39)	(.42)	(1.20)	(.05)
1990	6.38	6.65	80.86	7.18
	(.41)	(.44)	(1.25)	(.05)
2000	4.49	4.84	80.67	7.33
	(.54)	(.58)	(1.65)	(.07)
Change 1992 to 2000	−1.60	−1.59	−.88	.18
	(.47)	(.48)	(1.39)	(.07)
SER	.92	.93	.94	.91
TVP SE	.026	.023	.024	.028
(90% CI)	(.000 to .114)	(.000 to .106)	(.000 to .107)	(.000 to .115)

Notes: Results for the first colmn use the civilian unemployment rate as u in the estimated equation. The row labeled *Phillips curve slope* shows the estimates of β_π, with the standard error in parentheses and the p-value of the QLR stability test in brackets. *Trend values* are the (univariate) low-pass estimates of trend unemployment rate, and *NAI trend values* are the estimated values of the NAIRU computed using the Kalman smoother, as described in the text. *SER* is the standard error of the regression, and the *TVP SE* is the median-unbiased estimate of the standard deviation of the change in the equation's constant term. The equation was estimated using 4 lags of $\Delta\pi_t$ and 2 lags of Δu_t; the vector Z contained Gordon's (1982) wage- and price-control variable, 2 lags of the relative price of food and energy, and 2 lags of the exchange rate. The sample period is 1960:1 to 2000:2. Results in the remaining columns replace u with the demographically adjusted unemployment rate, capacity utilization, and building permits.

specifications. The estimated value of the slope coefficient in the unemployment specification is comparable to estimates obtained elsewhere in this literature using different sample periods and different series (see Staiger, Stock, and Watson 1997a, table 1), although those other estimates are approximately half the size of the coefficients in Gordon (1998, table 3).

Second, the TV-NAIRU is estimated to have fallen by approximately 1.6 percentage points from 1992 to 2000; the decline is only slightly less if it is measured using the demographically adjusted unemployment rate. In contrast, the NAIRU for capacity utilization and building permits is relatively stable; for example, the decline in the capacity-utilization TV-NAIRU is only two-thirds of the Kalman smoother standard error of the estimated decline.[3]

Third, for all practical purposes, the estimated NAIRUs are simply the univariate trends in the various activity measures. This can be seen by comparing the selected univariate trend values with the estimated TV-NAIRUs shown in table 1.2 or by comparing the sample paths of these values for the specification plotted in figure 1.4. The reason why the TV-NAIRU and trend values are so similar follows from the data analysis presented in the first section. Neglecting lags and the supply shocks, the Phillips curve is

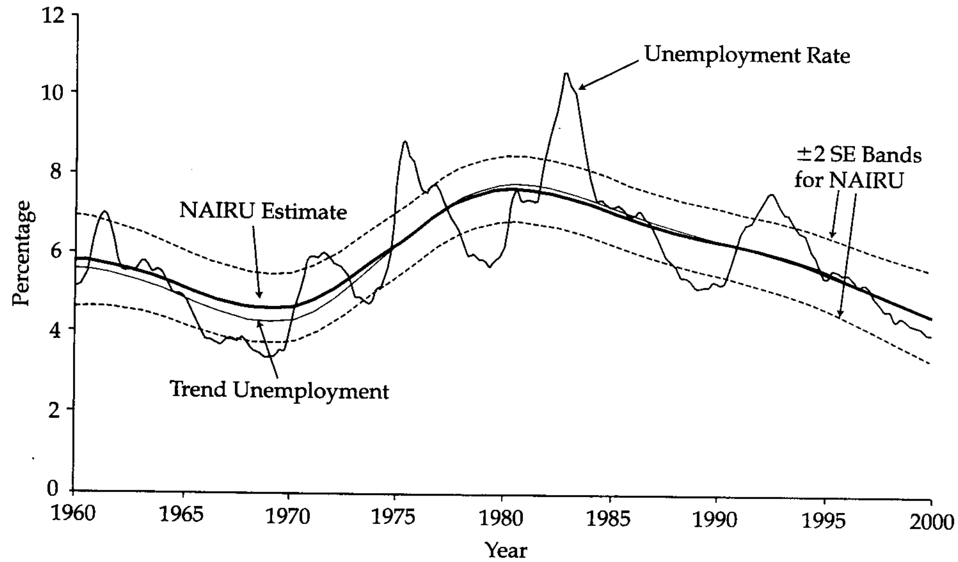
$$\Delta\pi_{t+1} = \beta_\pi(u_t^g) - \beta_\pi(u_t^N - u_t^*) + v_{\pi t+1}. \tag{1.13}$$

From the data analysis in the first section, $\Delta\pi_{t+1}$ is I(0), and, by construction, so is the unemployment-rate gap, u_t^g. This means that there cannot be large persistent deviations of u_t^N from u_t^*: if there were, these would be transmitted to $\Delta\pi$, but, since $\Delta\pi$ is I(0), it does not contain large persistent movements. Mechanically, this means that, in all the specifications, the median-unbiased estimate of the standard deviation of the intercept drift is very small; indeed, it is nearly the same value in each specification, between 0.023 and 0.028. This corresponds to a change in the intercept between 0.046 and 0.056 percentage points per year, which is nearly two orders of magnitude less than the standard deviation of the dependent variable, the quarterly change in inflation at an annual rate. In all the specifications, the 90 percent confidence interval includes 0, so the hypothesis of no parameter drift in these equations cannot be rejected at the 5 percent significance level.

While the TV-NAIRU and the univariate trend are very similar, figure 1.4 does show some differences between u_t^N and u_t^*. The deviation in the 1960s and early 1970s is associated with the trend increase in inflation over this period, and the deviation in the early 1980s is associated with a decline in the trend rate of inflation.

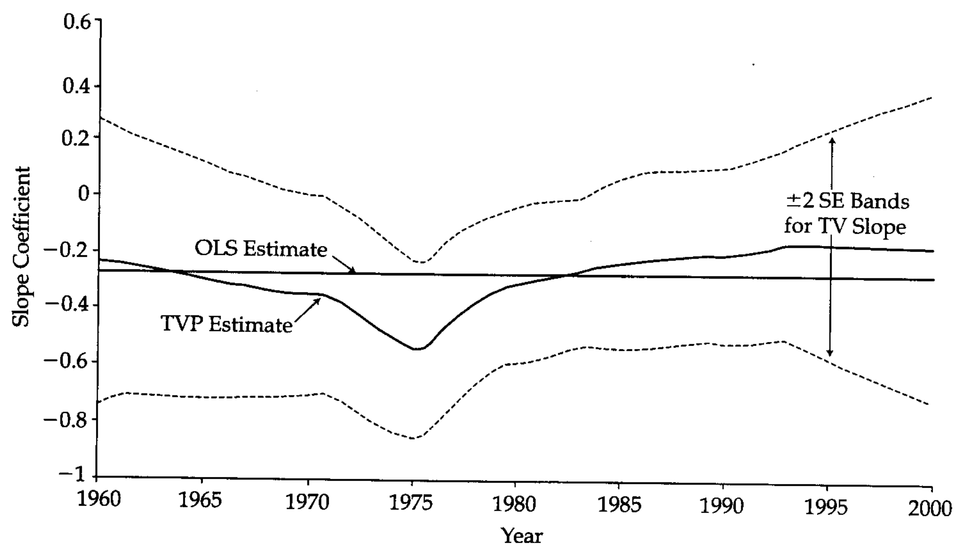
The fourth result from table 1.2 concerns the time variation in the slope of the Phillips curve. The p-values for tests of no slope change range from 0.02 to 0.16, suggesting possible time variation in the slope. To investigate the magnitude and timing of this variation, we estimated a model that allowed the slope coefficient to vary but held the intercept constant. Specifically, β_π was modeled as a random walk with innovation variance estimated using the method described in Stock and Watson (1998). Figure 1.5 shows the estimates of the time-varying slope

FIGURE 1.4 / The NAIRU from the Price Phillips Curve



Sources: Bureau of Labor Statistics and authors' calculations.

FIGURE 1.5 / Slope from the Baseline Price Phillips Curve



Source: Authors' calculations.

Notes: TV = time varying; TVP = time varying parameter.

coefficients for the specification using the unemployment rate obtained by the Kalman smoother together with ± 2 standard-error bands and the OLS slope estimate. Most of the time variation evident in the slope occurs in the mid-1970s, and the estimated slope remained essentially unchanged during the 1990s. Similar results obtain using the other variables (capacity utilization, building permits, and the demographically adjusted unemployment rate). These results are consistent with some small amount of time variation in the Phillips curve slope over the entire sample period but little time variation in the past dozen or so years.

Sensitivity Analysis

Table 1.3 summarizes thirty-six alternative Phillips curve regressions that examine the sensitivity of the benchmark results presented in table 1.2. These regressions differ by the price index used to measure inflation, the activity measure used, whether supply-shock control variables are included, whether the error-correction term and its lags are included, whether the log level of the productivity-adjusted real wage (the "markup") is included, and how many lags are included in the specifications. The statistics reported in the table are the same as those reported in table 1.2, except that, to save space, the values of the level of the trend activity measure and the associated TV-NAIRU are not reported; rather, only the change in the TV-NAIRU from 1992 to 2000 (and its Kalman smoother standard error, ignoring estimation uncertainty) is reported.

These results suggest eight conclusions.

First, the specifications in which the unemployment-rate gap is the activity variable are robust to these changes. The coefficient on the unemployment-rate gap is fairly stable, with estimates ranging from $-.25$ to $-.37$ across specifications and all the estimated coefficients within a standard error of -0.3. The TV-NAIRUs estimated with the unemployment-rate specifications are all estimated to have declined substantially from 1992 to 2000, with almost all the estimated declines being approximately 1.4 to 1.7 percentage points.

Second, in virtually all the specifications with alternative activity gaps, the activity-gap coefficients are significant at the 5 percent level (usually at the 1 percent level). Thus, the evidence is consistent with there being a generalized Phillips relation, where the unemployment rate is only one of several possible indicators that can be used in this relation.

Third, in virtually all the specifications, the median-unbiased estimator of the drift in the intercept term suggests that there is very little drift in the intercept in these regressions. In almost all specifications, the null hypothesis of no parameter drift is not rejected at the 5 percent significance level.

Fourth, there is also little evidence of substantial time variation in the slope of the Phillips curve. Only a few of the test statistics for time variation are significant at the 5 percent level (six of thirty-six), and, when time variation is allowed,

(Text continues on p. 26.)

TABLE 1.3 / Alternative Phillips Curve Estimates from Macroeconomic Data: Price Inflation Equation

$$\Delta\pi_{t+1} = \beta_\pi(u_t - u_t^N) + \alpha_{\pi\pi}(L)\Delta\pi_t + \alpha_{\pi x}(L)\Delta u_t + \alpha_{\omega\omega}(L)(\omega_t - \theta_{t-1}^* - \pi_{t-1}) + \alpha_{\omega\omega}(L)\Delta\omega_t + \gamma_\pi Z_t + \nu_{\pi t+1}$$

Price	Activity	Np	Na	Ng	Nr	Nx	Ne	Nw	Nm	PC Slope	SER	TVPSE	ΔNAITV: Median UB	ΔIntercept: Median UB	ΔNAITV: 5%	ΔIntercept: 5%
A. Baseline specification (table 1.1)																
GDP def.	Civ. unemp.	4	2	1	2	2	0	0	0	−.28 (.10) [.06]	.92	.026	−1.60 (.47)	−.00 (.13)	−1.62 (.68)	−.01 (.19)
GDP def.	Dem adj. UR	4	2	1	2	2	0	0	0	−.26 (.09) [.13]	.93	.023	−1.59 (.48)	−.00 (.12)	−1.62 (.75)	−.01 (.19)
GDP def.	Cap. util.	4	2	1	2	2	0	0	0	.09 (.03) [.02]	.94	.024	−.88 (1.39)	.03 (.13)	−1.19 (2.14)	.06 (.19)
GDP def.	Bldg. perm.	4	2	1	2	2	0	0	0	2.18 (.44) [.16]	.91	.028	.18 (.07)	−.02 (.14)	.20 (.09)	−.05 (.19)
B. Different price indices																
PCE def.	Civ. unemp.	4	2	1	2	2	0	0	0	−.28 (.11) [.20]	1.08	.017	−1.61 (.33)	−.00 (.09)	−1.59 (.71)	.00 (.20)
PCE def.	Dem adj. UR	4	2	1	2	2	0	0	0	−.26 (.11) [.29]	1.09	.015	−1.60 (.31)	−.00 (.08)	−1.58 (.75)	.00 (.20)
PCE def.	Cap. util.	4	2	1	2	2	0	0	0	.10 (.03) [.49]	1.07	.020	−.70 (1.08)	.01 (.11)	−1.05 (1.95)	.05 (.20)
PCE def.	Bldg. perm.	4	2	1	2	2	0	0	0	2.02 (.54) [.19]	1.08	.021	.18 (.05)	−.01 (.11)	.19 (.10)	−.03 (.20)
CPI	Civ. unemp.	4	2	1	2	2	0	0	0	−.39 (.14) [.16]	1.34	.000	−1.60 (.00)	.00 (.00)	−1.57 (.53)	.01 (.20)

CPI	Dem adj. UR	4	2	1	2	2	0	0	0	−.37 (.14) [.19]	1.35	.000	−1.59 (.00)	.00 (.00)	−1.56 (.55)	.01 (.20)
CPI	Cap. util.	4	2	1	2	2	0	0	0	.15 (.04) [.19]	1.34	.000	−.57 (.00)	.00 (.00)	−1.04 (1.40)	.07 (.20)
CPI	Bldg. perm.	4	2	1	2	2	0	0	0	3.63 (.71) [.05]	1.37	.028	.18 (.04)	−.01 (.15)	.18 (.06)	−.03 (.20)
C. Adding lags of wage inflation and error-correction term																
GDP def.	Civ. unemp.	4	2	1	2	2	1	4	0	−.25 (.10) [.07]	.91	.026	−1.67 (.55)	−.02 (.14)	−1.75 (.77)	−.04 (.19)
GDP def.	Dem adj. UR	4	2	1	2	2	1	4	0	−.23 (.10) [.16]	.92	.024	−1.65 (.56)	−.01 (.13)	−1.76 (.84)	−.04 (.19)
GDP def.	Cap. util	4	2	1	2	2	1	4	0	.09 (.03) [.02]	.94	.025	−.74 (1.49)	.02 (.13)	−.83 (2.21)	.02 (.19)
GDP def.	Bldg. perm.	4	2	1	2	2	1	4	0	2.05 (.44) [.20]	.91	.030	.20 (.07)	−.04 (.15)	.21 (.09)	−.07 (.19)
D. Increasing lag length																
GDP def.	Civ. unemp.	8	4	1	2	2	0	0	0	−.34 (.12) [.34]	.89	.052*	−1.71 (.68)	−.04 (.23)	−1.66 (.55)	−.02 (.19)
GDP def.	Dem adj. UR	8	4	1	2	2	0	0	0	−.31 (.12) [.27]	.91	.050*	−1.70 (.74)	−.03 (.23)	−1.66 (.62)	−.02 (.19)
GDP def.	Cap. util.	8	4	1	2	2	0	0	0	.10 (.03) [.00]	.94	.043	−1.15 (1.95)	.06 (.20)	−1.10 (1.83)	.06 (.19)
GDP def.	Bldg. perm.	8	4	1	2	2	0	0	0	2.29 (.50) [.23]	.92	.042	.19 (.09)	−.05 (.20)	.19 (.08)	−.04 (.19)

(Table continues on p. 24.)

TABLE 1.3 / *Continued*

Price	Activity	Np	Na	Ng	Nr	Nx	Ne	Nw	Nm	PC Slope	SER	TVPSE	ΔNAITV: Median UB	ΔIntercept: Median UB	ΔNAITV: 5%	ΔIntercept: 5%
GDP def.	Civ. unemp.	8	4	1	4	4	0	0	0	−.34 (.12) [.42]	.89	.044*	−1.67 (.60)	−.02 (.21)	−1.65 (.56)	−.02 (.19)
GDP def.	Dem adj. UR	8	4	1	4	4	0	0	0	−.31 (.12) [.33]	.91	.041	−1.65 (.65)	−.02 (.20)	−1.65 (.62)	−.02 (.19)
DGP def.	Cap. util.	8	4	1	4	4	0	0	0	.11 (.03) [.01]	.94	.036	−1.06 (1.69)	.05 (.18)	−1.11 (1.81)	.06 (.19)
GDP def.	Bldg. perm.	8	4	1	4	4	0	0	0	2.25 (.51) [.38]	.92	.043	.20 (.09)	−.05 (.20)	.19 (.09)	−.04 (.19)
E. Eliminating the relative price of food and energy, exchange rates, and wage- and price-control variables																
GDP def.	Civ. unemp.	4	2	0	0	0	0	0	0	−.37 (.10) [.08]	1.00	.048	−1.55 (.61)	.02 (.23)	−1.56 (.53)	.01 (.19)
GDP def.	Dem adj. UR	4	2	0	0	0	0	0	0	−.34 (.10) [.17]	1.02	.048	−1.53 (.67)	.02 (.23)	−1.55 (.57)	.01 (.20)
GDP def.	Cap. util.	4	2	0	0	0	0	0	0	.11 (.03) [.05]	1.02	.050*	−1.40 (2.09)	.09 (.23)	−1.17 (1.74)	.07 (.20)
GDP def.	Bldg. perm.	4	2	0	0	0	0	0	0	1.78 (.44) [.24]	1.03	.031	.18 (.09)	−.01 (.16)	.18 (.11)	−.01 (.20)
F. Adding markup variable																
GDP def.	Civ. unemp.	4	2	1	2	2	0	0	2	−.27 (.09) [.09]	.91	.035	−1.43 (.65)	.05 (.17)	−1.41 (.72)	.05 (.19)

GDP def.	Dem adj. UR	4	2	1	2	2	0	0	2	−.24 (.09) [.19]	.92	.035	−1.40 (.73)	.05 (.17)	−1.38 (.80)	.05 (.19)
GDP def.	Cap. util.	4	2	1	2	2	0	0	2	.09 (.03) [.02]	.93	.023	−.96 (1.41)	.03 (.12)	−1.36 (2.24)	.07 (.19)
GDP def.	Bldg. perm.	4	2	1	2	2	0	0	2	2.08 (.44) [.26]	.90	.025	.16 (.06)	.02 (.13)	.16 (.09)	.03 (.19)
GDP def.	Civ. unemp.	8	4	1	2	2	0	0	2	−.32 (.12) [.34]	.89	.043*	−1.44 (.64)	.05 (.20)	−1.45 (.60)	.05 (.19)
GDP def.	Dem adj. UR	8	4	1	2	2	0	0	2	−.28 (.12) [.27]	.90	.040	−1.43 (.69)	.05 (.19)	−1.43 (.68)	.05 (.19)
GDP def.	Cap. util.	8	4	1	2	2	0	0	2	.10 (.03) [.01]	.93	.036	−1.18 (1.80)	.06 (.18)	−1.25 (1.94)	.07 (.19)
GDP def.	Bldg. perm.	8	4	1	2	2	0	0	2	2.20 (.50) [.28]	.91	.032	.16 (.07)	.03 (.16)	.16 (.09)	.03 (.19)

Notes: This table contains alternative estimates of the price Phillips curve (PC) equation and the TV-NAIRU. The stub column shows the price series used to construct π, and the first column shows the activity variable used for u. the next columns show a set of parameters that described the specification: Np = number of lags of inflation; Na = number of lags of activity variable; NG = a binary variable indicating inclusion(1)/exclusion(0) of wage and price variable (Gordon 1982); Nr = number of lags of the relative price of food and energy; Nx = number of lags of exchange rates; Ne = a binary variable indicating inclusion(1)/exclusion(0) of the wage/price/productivity error-correction term $(\omega_t - \theta^*_{t-1} - \pi_{t-1})$; Nw = number of lags of $\Delta\omega$; and Nm = number of lags of markup variable described in the text. The next three columns are described in the notes to table 1.1. An asterisk in the column labeled *TVPSE* indicates that the estimated TVPSE is significantly different from zero at the 5 percent level. the columns labeled Δ*NAITV* show the change in the NAIRU (or the NAI trend value) from 1992:1 to 2000:1, and the columns labeled Δ*Intercept* shows the implied in the equation's constant term over the same period. These are presented for two values of the standard error of the change in the constant: the median-unbiased estimate and a value of .039 (which implies a standard deviation of the change in the constant of 5 percent over the sample period). The sample period is 1960:1 to 2000:1 for all estimated equations.

the estimated sample path of the time-varying slope shows little movement over the past decade.[4]

Fifth, the estimate of the 1992 to 2000 change in the TV-NAIRU is largely unaffected by how supply shocks are treated. For example, in the benchmark specification (which includes the supply-shock variables), the NAIRU is estimated to decline by 1.60 percentage points, whereas, if the supply-shock variables are omitted, the NAIRU is estimated to decline by 1.55 percentage points. However, the regression standard errors of the specifications with supply shocks included are significantly smaller than those with the supply shock omitted. Evidently, these variables are important for explaining one-off changes in inflation but not the kind of persistent changes that could be confused with a change in the NAIRU.

Sixth, the estimated recent decline in the TV-NAIRU is essentially unaffected by whether the total unemployment rate or the demographically adjusted unemployment rate is used. This is consistent with the discussion in Gordon (1998) and Stock (in press) concluding that, although demographic shifts might be associated with increases in the NAIRU in the 1970s, the timing of demographic shifts is not aligned with this sharp recent decline.

Seventh, these results confirm the finding in columns 3 and 4 of table 1.2 that TV-NAIRUs estimated using the rate of capacity utilization and building permits have been relatively stable; for capacity utilization, the change from 1992 to 2000 is less than its Kalman smoother standard error.

Eighth, adding the error-correction term to the benchmark specification decreases the standard error of the regression slightly but does not change the estimates of the slope coefficient of the TV-NAIRU. This suggests that the estimated decline in the NAIRU is not a spurious consequence of neglecting feedback from wages to prices. Table 1.3 also shows results for a specification in which the markup of prices over productivity-adjusted wages (or, equivalently, the log level of the productivity-adjusted real wage) is included. Including this variable reduces somewhat the estimated decline of the unemployment NAIRU, from 1.60 percentage points in the base specification to 1.43 in the specification including this term. Thus, this term is estimated to contribute perhaps 0.2 percentage points to the decline in the NAIRU. Taken together, these results suggest that there is limited or no evidence that feedback from wages to prices has served to hold down prices during the 1990s.

Summary of Main Findings

The regression results reported in tables 1.2 and 1.3 indicate a stable and statistically significant relation between future changes in price inflation and current economic activity as measured by various activity gaps. In addition, in these gap specifications, there is very little evidence of drift in the intercept or the slope, either in terms of statistical significance or in terms of the point estimates of the drift from 1992 to 2000. Finally, including supply shocks does not change substantially the estimates of the declines in the NAIRU.

We interpret these findings as being inconsistent with "the Phillips curve is dead" theories of the 1990s. They are inconsistent with theories that place considerable weight on changes in price-setting behavior in the 1990s, for these theories would imply important drift in the intercept or slope of the Phillips curve. They are also inconsistent with theories that place great weight on sustained "good luck" in the form of favorable supply shocks. Said differently, once the Phillips curves are specified in gaps, there are no price-equation puzzles to explain. Because trend capacity utilization and trend building permits are approximately flat, the only "puzzle" about the price Phillips curve is why the univariate trend in the unemployment rate has fallen. Once we have accounted for the univariate trend in the unemployment rate, these price-inflation Phillips curves fit quite nicely throughout the decade and, indeed, throughout the entire sample period, 1960 to 2000.

MACRO ESTIMATES OF WAGE PHILLIPS CURVES

This section presents empirical estimates of wage Phillips curves and TV-NAIRUs using the unemployment rate and other indicators of economic activity. The discussion parallels that presented in the previous section: first, we present some benchmark estimates; next, we examine the robustness of these estimates to alternative specifications; and, finally, we summarize conclusions.

Benchmark Wage Regressions

BENCHMARK REGRESSION ESTIMATES Benchmark wage regressions are reported in table 1.4, using the same format as in table 1.2.

The most striking result to be seen in table 1.4 is that these specifications are very similar to the benchmark price regressions reported in the corresponding columns of table 1.2. The slope coefficients are larger in table 1.4 than in table 1.2, but so is the standard deviation of the dependent variable. The slope coefficients are all statistically significant at the 5 percent level. Although the levels of the TV-NAIRUs are different in table 1.2 than in table 1.4 (because the variables have different means), the *changes* in the TV-NAIRUs are almost the same.[5] For example, on the basis of the total unemployment rate in the price equation in column 1 of table 1.2, the TV-NAIRU is estimated to decline by 1.60 percentage points from 1992 to 2000; on the basis of the labor-share specification in table 1.4, this decline is estimated to be 1.52 percentage points. The quantitative declines in the TV-NAIRUs are the same for the other activity variables in the two tables.

A key similarity between the price results in table 1.2 and the wage results in table 1.4 is that the intercept drift is negligible in both tables. Although the me-

TABLE 1.4 / Phillips Curve Estimates from Macroeconomic Data: Wage Inflation Equation $\omega_{t+1} - \theta_t^* - \pi_t = \beta_\omega(u_t - u_t^N) + \alpha_{\omega\pi}(L)\Delta\pi_t + \alpha_{\omega\omega}(L)(\omega_t - \theta_{t-1}^* - \pi_{t-1}) + \alpha_{\omega u}(L)\Delta u_t + \nu_{\omega t+1}$

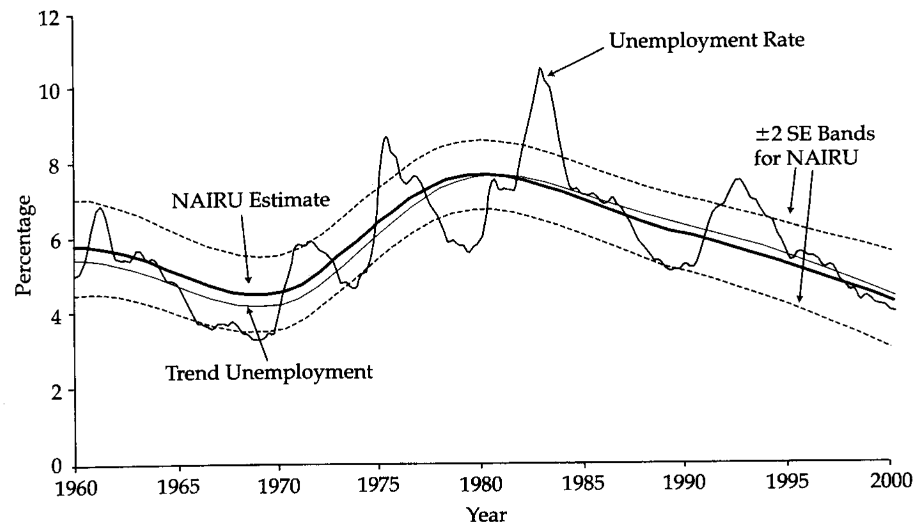
	Civilian Unemployment Rate	Demographically Adjusted Unemployment Rate	Capacity Utilization	Building Permits
Phillips Curve Slope (SE) [stability p-value]	−.42 (.19) [.10]	−.39 (.18) [.10]	.11 (.05) [.29]	2.00 (.78) [.99]
Trend values				
1970	4.33	4.39	85.51	7.27
1980	7.79	7.55	78.74	7.23
1990	6.40	6.66	80.84	7.18
2000	4.48	4.83	81.02	7.31
Change 1992 to 2000	−1.60	−1.59	−.57	.17
NAI trend values				
1970	3.75 (.46)	3.72 (.49)	88.61 (1.37)	7.40 (.10)
1980	6.98 (.43)	6.67 (.46)	81.92 (1.34)	7.41 (.09)
1990	5.39 (.46)	5.58 (.49)	84.09 (1.37)	7.40 (.10)
2000	3.54 (.58)	3.80 (.61)	84.26 (1.43)	7.52 (.12)
Change 1992 to 2000	−1.52 (.43)	−1.52 (.43)	−.59 (.45)	.16 (.09)
SER	1.81	1.81	1.82	1.81
TVP SE (90 percent CI)	.034 (.000 to .186)	.031 (.000 to .182)	.009 (.000 to .150)	.034 (.000 to .187)

Notes: The equation was estimated using four lags of $\Delta\pi_t$, four lags of $(\omega_t - \theta_{t-1}^* - \pi_{t-1})$, and two lags of Δu_t. For a description of the table entries, see the note to table 1.1.

dian-unbiased estimate is larger in table 1.4 than in table 1.2, the dependent variable in table 1.4 is more variable, and the standard error of the wage regressions is twice that of the price regressions, so the relative variability in the intercept is virtually identical in the price and wage specifications. In all specifications in table 1.4, the 90 percent confidence interval for the standard deviation of the change in the intercept includes 0; that is, the hypothesis of no parameter drift cannot be rejected in any of these specifications at the 5 percent significance level.

Figure 1.6 plots the estimated TV-NAIRU for unemployment on the basis of the specification in column 1 of table 1.4, where the TV-NAIRU is adjusted so

FIGURE 1.6 / The NAIRU from the Wage Phillips Curve



Sources: Bureau of Labor Statistics and authors' calculations.

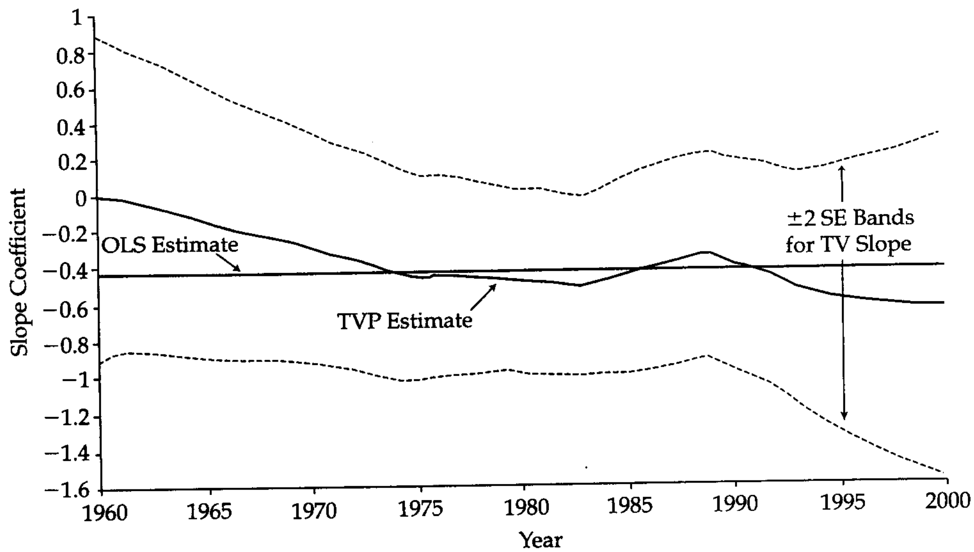
that it has the same sample mean as the univariate trend in the unemployment rate (see note 5). Inspection of this figure underscores that there is effectively no difference between the TV-NAIRU and the univariate trend in unemployment. This is the same conclusion as was drawn from the price TV-NAIRU plotted in figure 1.4. Comparison of figures 1.4 and 1.6 reveals that the TV-NAIRUs estimated from the price Phillips curve in column 1 of table 1.2 and the wage Phillips curve in column 1 of table 1.4 (shifted per note 5) are essentially identical. The reason for this is that, in both specifications, the median-unbiased estimate indicates negligible intercept drift.

Table 1.4 indicates that there is little evidence of changes in the slope of the wage Phillips curve. The *p*-values for the test of the null hypothesis of no change in the slope range from 0.10 to 0.99. Figure 1.7 shows the estimated values of the time-varying slope in the unemployment-rate wage Phillips curve using a specification that parallels the results for the price Phillips curve shown in figure 1.5. The point estimates suggest a slight steepening of the wage Phillips curve over the past decade (consistent with the scatterplot in figure 1.2), but the standard-error bands make it clear that these changes are far from statistically significant.

Sensitivity Analysis

Table 1.5 summarizes results for forty-eight variations of the wage Phillips curve. Some of these variations are similar to the sensitivity checks reported in

FIGURE 1.7 / Slope from the Baseline Wage Phillips Curve



Source: Authors' calculations.

Notes: TV = time varying; TVP = time varying parameter.

table 1.3, for example, changing the definition of the wage series, changing the number of lags, and so on. Inspection of table 1.5 reveals that the main conclusions to be drawn from table 1.4, particularly the lack of intercept and slope drift in the gap specifications, are robust to these changes. However, there are some important differences among these specifications.

Most notably, these wage specifications are less stable than the price specifications are to changes in definitions of the variables. For example, the slope coefficient is often insignificant in specifications in which the GDP deflator is replaced by the PCE deflator or the CPI as well as in specifications in which wage growth is measured using the ECI (total compensation or wages and salaries only) or average hourly earnings. (The sample period for the specifications using the ECI data was limited to 1982 to 1999.) The results using compensation per hour are, however, consistent with the benchmark results. Overall, however, these results suggest that the estimated Phillips curve using trend unit labor costs is rather delicate.

An important result to be seen in table 1.5 is that replacing trend productivity growth with its sample average growth rate results in coefficients on the activity variable that are essentially unchanged but induces intercept drift that is both economically large and, now, statistically significant. In contrast to the bench-

(Text continues on p. 36.)

TABLE 1.5 / Alternative Phillips Curve Estimates from Macroeconomic Data: Wage Inflation Equation

$$\omega_{t+1} - \theta_t^* - \pi_t = \beta_\omega (u_t - u_t^N) + \alpha_{\omega\pi}(L)\Delta\pi_t + \alpha_{\omega\omega}(L)(\omega_t - \theta_{t-1}^* - \pi_{t-1}) + \alpha_{\omega u}(L)\Delta u_t + \nu_{\omega t+1}$$

Wage	Price	Activity	Nw	Na	Np	Pr	PC Slope	SER	TVPSE	Δ NAITV: Median UB	Δ Intercept: Median UB	Δ NAITV: 5 percent	Δ Intercept: 5 percent
A. Baseline specification (table 1.3)													
Comp./hr	GDP Def.	Civ. unemp.	4	2	4	2	−.42 (.19) [.10]	1.81	.034	−1.52 (.43)	.03 (.18)	−1.49 (.50)	.05 (.21)
Comp./hr	GDP Def.	Dem adj. UR	4	2	4	2	−.39 (.18) [.10]	1.81	.031	−1.52 (.43)	.03 (.17)	−1.47 (.54)	.05 (.21)
Comp./hr	GDP Def.	Cap. util.	4	2	4	2	.11 (.05) [.29]	1.82	.009	−.59 (.45)	.00 (.05)	−1.18 (1.93)	.07 (.21)
Comp./hr	GDP Def.	Bldg. perm.	4	2	4	2	2.01 (.78) [.99]	1.81	.034	.16 (.09)	.03 (.19)	.15 (.10)	.04 (.21)
B. Alternative productivity trends													
Comp./hr	GDP Def.	Civ. unemp.	4	2	4	1	−.42 (.19) [.14]	1.86	.102*	−.76 (1.10)	.35 (.47)	−1.42 (.50)	.07 (.21)
Comp./hr	GDP Def.	Dem adj. UR	4	2	4	1	−.38 (.19) [.13]	1.86	.101*	−.67 (1.22)	.35 (.46)	−1.40 (.55)	.07 (.21)
Comp./hr	GDP Def.	Cap. util.	4	2	4	1	.12 (.05) [.56]	1.87	.090*	−3.73 (3.59)	.37 (.42)	−1.36 (1.79)	.09 (.21)
Comp./hr	GDP Def.	Bldg. perm.	4	2	4	1	1.79 (.81) [.84]	1.87	.076	.04 (.21)	.23 (.37)	.13 (.12)	.07 (.21)
Comp./hr	GDP Def.	Civ. unemp.	4	2	4	3	−.40 (.19) [.33]	1.80	.056	−1.57 (.73)	.01 (.29)	−1.61 (.53)	−.00 (.21)

(Table continues on p. 32.)

TABLE 1.5 / *Continued*

Wage	Price	Activity	Nw	Na	Np	Pr	PC Slope	SER	TVPSE	Δ NAITV: Median UB	Δ Intercept: Median UB	Δ NAITV: 5 percent	Δ Intercept: 5 percent
Comp./hr	GDP Def.	Dem adj. UR	4	2	4	3	−.36 (.18) [.32]	1.80	.056	−1.56 (.79)	.01 (.29)	−1.60 (.58)	−.00 (.21)
Comp./hr	GDP Def.	Cap. util.	4	2	4	3	.11 (.05) [.21]	1.81	.046	−.81 (2.21)	.03 (.24)	−.72 (1.94)	.02 (.21)
Comp./hr	GDP Def.	Bldg. perm.	4	2	4	3	1.84 (.78) [.75]	1.81	.058	.17 (.16)	−.00 (.30)	.18 (.11)	−.01 (.21)
C. Alternative price indices													
Comp./hr	NFB Def.	Civ. unemp.	4	2	4	2	−.36 (.19) [.24]	1.87	.025	−1.58 (.38)	.01 (.14)	−1.52 (.59)	.03 (.21)
Comp./hr	NFB Def.	Dem adj. UR	4	2	4	2	−.32 (.18) [.24]	1.87	.015	−1.58 (.27)	.00 (.09)	−1.50 (.66)	.03 (.21)
Comp./hr	NFB Def.	Cap. util	4	2	4	2	.11 (.05) [.43]	1.88	.000	−.57 (.00)	.00 (.00)	−1.08 (2.00)	.05 (.21)
Comp./hr	NFB Def.	Bldg. perm.	4	2	4	2	2.94 (.84) [.78]	1.84	.017	.17 (.03)	.00 (.09)	.17 (.07)	.02 (.21)
Comp./hr	PCE Def.	Civ. unemp.	4	2	4	2	−.29 (.19) [.10]	1.84	.088*	−.78 (1.42)	.24 (.42)	−1.43 (.72)	.05 (.21)
Comp./hr	PCE Def.	Dem adj. UR	4	2	4	2	−.26 (.18) [.10]	1.84	.086	−.71 (1.56)	.23 (.41)	−1.40 (.81)	.05 (.21)
Comp./hr	PCE Def.	Cap. util.	4	2	4	2	.08 (.05) [.50]	1.84	.074	−3.61 (4.72)	.23 (.36)	−1.45 (2.73)	.07 (.21)

Comp./hr	PCE Def.	Bldg. perm.	4	2	4	2	1.73 (.83) [.89]	1.83	.087	.05 (.24)	.22 (.41)	.15 (.12)	.05 (.21)
Comp./hr	CPI	Civ. unemp.	4	2	4	2	−.17 (.20) [.13]	1.85	.098*	−.86 (2.59)	.13 (.45)	−1.53 (1.21)	.01 (.21)
Comp./hr	CPI	Dem adj. UR	4	2	4	2	−.15 (.19) [.12]	1.85	.095*	−.78 (2.97)	.12 (.44)	−1.51 (1.41)	.01 (.21)
Comp./hr	CPI	Cap. util.	4	2	4	2	.06 (.06) [.71]	1.86	.085	−3.27 (6.83)	.16 (.41)	−1.09 (3.54)	.03 (.21)
Comp./hr	CPI	Bldg. perm.	4	2	4	2	2.04 (.87) [.78]	1.83	.109*	.13 (.24)	.08 (.49)	.17 (.10)	−.00 (.21)
D. Alternative wage indices													
AHE PW	GDP Def.	Civ. unemp.	4	2	4	5	−.15 (.12) [.55]	1.12	.077*	−1.33 (2.18)	.04 (.33)	−1.46 (1.32)	.02 (.20)
AHE PW	GDP Def.	Dem adj. UR	4	2	4	5	−.14 (.12) [.48]	1.13	.077*	−1.29 (2.31)	.04 (.33)	−1.44 (1.39)	.02 (.20)
AHE PW	GDP Def.	Cap. util.	4	2	4	5	.04 (.03) [.35]	1.11	.061*	−1.89 (6.85)	.05 (.28)	−1.33 (4.86)	.03 (.20)
AHE PW	GDP Def.	Bldg. perm.	4	2	4	5	1.08 (.55) [.70]	1.11	.086*	.17 (.33)	.00 (.35)	.17 (.18)	.00 (.20)
Comp./hr man.	GDP Def.	Civ. unemp.	4	2	4	5	−.50 (.21) [.45]	2.01	.141*	−1.69 (1.20)	−.05 (.60)	−1.71 (.43)	−.05 (.21)
Comp./hr man.	GDP Def.	Dem adj. UR	4	2	4	5	−.47 (.21) [.45]	2.02	.140*	−1.69 (1.27)	−.05 (.60)	−1.70 (.45)	−.05 (.21)

(Table continues on p. 34.)

TABLE 1.5 / *Continued*

Wage	Price	Activity	Nw	Na	Np	Pr	PC Slope	SER	TVPSE	Δ NAITV: Median UB	Δ Intercept: Median UB	Δ NAITV: 5 percent	Δ Intercept: 5 percent
Comp./hr man.	GDP Def.	Cap. util.	4	2	4	5	.13 (.06) [.44]	2.04	.137*	−1.19 (4.57)	.08 (.59)	−.29 (1.65)	−.04 (.21)
Comp./hr man.	GDP Def.	Bldg. perm.	4	2	4	5	.98 (.92) [.92]	2.06	.146*	.14 (.62)	.03 (.62)	.22 (.22)	−.04 (.21)
ECC-C	GDP Def.	Civ. unemp.	4	2	4	2	−.14 (.11) [.33]	.58	.000	−1.60 (.00)	.00 (.00)	−1.27 (1.22)	.05 (.17)
ECC-C	GDP Def.	Dem adj. UR	4	2	4	2	−.12 (.10) [.37]	.58	.000	−1.59 (.00)	.00 (.00)	−1.24 (1.37)	.04 (.17)
ECC-C	GDP Def.	Cap. util.	4	2	4	2	.03 (.04) [.01]	.59	.000	−.57 (.00)	.00 (.00)	−4.38 (6.30)	.10 (.17)
ECC-C	GDP Def.	Bldg. perm.	4	2	4	2	.99 (.55) [.66]	.58	.000	.17 (.00)	.00 (.00)	.14 (.17)	.04 (.17)
ECC-WS	GDP Def.	Civ. unemp.	4	2	4	2	−.06 (.12) [.10]	.76	.106*	5.16 (5.83)	.39 (.34)	1.52 (3.17)	.18 (.18)
ECC-WS	GDP Def.	Dem adj. UR	4	2	4	2	−.07 (.12) [.15]	.76	.104*	4.35 (5.06)	.39 (.33)	1.18 (2.78)	.18 (.18)
ECC-WS	GDP Def.	Cap. util.	4	2	4	2	.03 (.05) [.82]	.77	.088*	−10.64 (9.08)	.34 (.31)	−5.62 (5.43)	.17 (.18)
ECC-WS	GDP Def.	Bldg. perm.	4	2	4	2	2.02 (.72) [.10]	.73	.130*	−.06 (.18)	.47 (.36)	.06 (.09)	.22 (.18)

E. Increasing lag length													
Comp./hr	GDP Def.	Civ. unemp.	8	4	8	2	−.49 (.25) [.05]	1.79	.055	−1.37 (.57)	.11 (.28)	−1.48 (.43)	.06 (.21)
Comp./hr	GDP Def.	Dem adj. UR	8	4	8	2	−.45 (.24) [.06]	1.81	.053	−1.35 (.62)	.10 (.28)	−1.46 (.47)	.06 (.21)
Comp./hr	GDP Def.	Cap. util.	8	4	8	2	.12 (.07) [.33]	1.81	.029	−.91 (1.29)	.04 (.16)	−1.19 (1.70)	.08 (.21)
Comp./hr	GDP Def.	Bldg. perm.	8	4	8	2	2.47 (.92) [.79]	1.82	.027	.16 (.06)	.02 (.15)	.15 (.09)	.05 (.21)
F. Eliminating lags of price inflation													
Comp./hr	GDP Def.	Civ. unemp.	4	2	0	2	−.12 (.17) [.01]	1.93	.000	−1.60 (.00)	.00 (.00)	−1.16 (1.72)	.05 (.21)
Comp./hr	GDP Def.	Dem adj. UR	4	2	0	2	−.09 (.17) [.01]	1.93	.000	−1.59 (.00)	.00 (.00)	−.97 (2.46)	.05 (.21)
Comp./hr	GDP Def.	Cap. util.	4	2	0	2	.01 (.05) [.22]	1.93	.000	−.57 (.00)	.00 (.00)	−6.79 (23.06)	.06 (.21)
Comp./hr	GDP Def.	Bldg. perm.	4	2	0	2	1.07 (.78) [.85]	1.91	.000	.17 (.00)	.00 (.00)	.13 (.20)	.04 (.21)

Notes: This table contains alternative estimates of the wage Phillips curve (PC) equation and the TV-NAIRU. The stub column shows the wage series used to construct ω, the next column shows the price series used to construct π, and the second column shows the activity variable used for u. The next columns show a set of parameters that described the specification: Nw = number of lags of real wages; Np = number of lags of inflation; Na = number of lags of activity variable; Pr = trend estimate of productivity growth (1 = average value of nonfarm-business productivity growth, 1960 to 2000; 2 = low-pass filter of nonfarm-business productivity growth using I[0] AR extrapolation; 3 = low-pass filter of nonfarm-business productivity growth using I[1] AR extrapolation; 5 = low-pass filter of manufacturing productivity growth using I[0] AR extrapolation). The remaining columns are described in the notes to table 1.2. The sample period is 1960:1 to 2000:1 for all estimated equations except those including ECC-C and ECC-WS, which used the sample period 1982:1 to 2000:1. For descriptions of AHE PW, ECC-C, and ECC-WS, see the appendix. NFB = nonfarm business; PCE = personal consumption expenditure.

mark estimate, in which the unemployment-rate TV-NAIRU is estimated to have fallen by 1.52 percentage points from 1992 to 2000, the specification using average productivity growth shows a decline of only 0.76 percentage points in the estimated TV-NAIRU. Said differently, when the recent increase in trend productivity is excluded from the specification, the intercept in the Phillips curve adjusts to track the increase in real wages. The amount of the required adjustment is large: the change in the intercept implies an increase in the long-run mean change in the growth rate of wages of 0.84 percentage points.

Summary of Main Findings

These regressions point to a stable Phillips relation between trend unit labor costs and the various activity measures over this period, although the macro wage specifications are more delicate than are the macro price specifications examined in the third section. When the slope coefficient is precisely estimated, these specifications produce estimates of the 1992 to 2000 change in the TV-NAIRU that are strikingly similar to those produced by the price Phillips curves. In contrast, if the dependent variable is future wage inflation less current price inflation and the role of productivity growth is therefore ignored, then the behavior of the wage and price regressions is inconsistent, with real-wage inflation appearing in the second half of the 1990s. Specifications of the wage Phillips curve that ignore productivity growth appear unstable in the 1990s, while those that incorporate productivity growth are stable.

SPECIFICATION AND ESTIMATION OF STATE WAGE PHILLIPS CURVES

Specification of State-Level Wage Phillips Curves

The state regression specifications have the same basic form as the macro regressions, but data limitations lead to several modifications. For example, because the state data are annual, the timing conventions of the quarterly and annual specifications differ. Temporal aggregation—averaging both sides of equation 1.4 over the four quarters in the year—results in a relation between time series with dates that overlap by three quarters. We approximate this by using as the dependent variable $\omega_{t+1} - \theta_{t+1} - \pi_{t+1}$ (robustness to different timing conventions is investigated in the seventh section). Also, we use the unemployment rate as the activity variable in all the state Phillips curves.

These considerations lead to the state-level variant of equation 1.4,

$$\omega_{it+1} - \theta_{it+1} - \pi_{t+1} = \beta(u_{it} - u_{it}^{N}) + \zeta_t + \nu_{it+1}, \qquad (1.14)$$

where ω_{it+1} is the percentage growth in the nominal wage in state i from year t to year $t + 1$, θ_{it+1} is the annual percentage growth in labor productivity, u_{it} is the unemployment rate, u_{it}^N is the NAIRU for state i in year t (that is, the state TV-NAIRU), ζ_t are macro shocks (the sum of $\gamma_\omega Z_t$ and v_{t+1} in equation 1.4), and v_{it} is an error term that has mean 0 and that is uncorrelated with the macro shocks ζ_t. The subscript t runs over all the years in the sample, which differ slightly across specifications depending on data availability. As is discussed in the next section, in our data set, state nominal-wage growth ω_{it+1} and the unemployment rate u_{it} are computed from the Current Population Survey (CPS). State productivity θ_{it} is constructed in two different ways: either as the annual percentage growth of gross state product, less the growth of state employment, or from national industry-level productivity data weighted by the output share of each industry in the state.

The state TV-NAIRU can usefully be thought of as consisting of several components: the national TV-NAIRU (u_t^N); features that are unique to each state and constant over the sample, such as climate (ϕ_i); and institutional considerations that affect search and matching in the labor market, some of which are measured (X_{it}) and some of which are not (ε_{it}). That is, the state TV-NAIRU can be expressed as,

$$u_{it}^N = u_t^N + \phi_i + \gamma X_{it} + \varepsilon_{it}. \tag{1.15}$$

Substituting equation 1.15 into 1.14 and rearranging yields our base state-regression specification:

$$\omega_{it+1} - \theta_{it+1} - \pi_{t+1} = \alpha_i + \delta_t + \beta u_{it} + \gamma X_{it} + \nu_{it+1}, \tag{1.16}$$

where $\alpha_i = -\beta\phi_i$ are state effects, and $\delta_t = \zeta_t - \beta u_t^N$ are time effects.

It is worth making three remarks about the specification equation 1.16. First, unlike the macro regressions, this benchmark specification for the state panel regressions does not include lags of either the unemployment rate or the labor share. The reason for this is practical: with only twenty annual observations, it is unlikely that we will estimate lag dynamics with any precision, and, in any event, the lag dynamics will be less pronounced at the annual level than at the quarterly level used in the macro data. In sensitivity checks, however, we report the results of specifications that include lags.

Second, as we have already discussed, there is some debate over whether the correct specification of this model should include a lagged wage level on the right-hand side—that is, should the model be specified in terms of wage levels (the "wage curve") or real-wage growth less productivity (the Phillips curve)? Specifications using wage growth have the implication that states' productivity-adjusted real wages can drift arbitrarily far apart over long periods, and this is implausible since capital and labor can flow across state boundaries. However, the empirical evidence suggests that capital and labor migrate slowly enough that the Phillips curve specification fits that data better than wage-curve specifications with substantial mean reversion (see Blanchard and Katz 1992, 1997; Card and Hyslop 1997; and Autor and Staiger 2001). This leads us to use the Phillips curve as our benchmark specification, although, in our sensitivity analysis, we consider specifications that include the levels of productivity-adjusted real wages.

Third, π_t is not indexed by state in equation 1.16. This is because data on prices by state are not available; thus, the deflation process uses the national price level (the Consumer Price Index for All Urban Consumers [CPI-U]). Because equation 1.16 includes year effects, the estimates of the slope coefficients β and γ are invariant to which inflation variable is used (the CPI, the GDP deflator, and so on), whether the inflation variable used is dated $t + 1$ or t, or, indeed, whether the dependent variable is deflated at all (that is, is $\omega_{it+1} - \theta_{it+1}$). The deflator is used only to identify the time effects $\{\delta_t\}$ and, thereby, to identify the macro TV-NAIRU from this state specification.

Estimation of a National TV-NAIRU from the State Regressions

Estimates of the annual national TV-NAIRU can be obtained from the year effects in the state regressions. The year fixed effects contain movements in the national TV-NAIRU, macro shocks, and estimation error. Thus, these year fixed effects must be filtered to obtain estimates of the national TV-NAIRU.

The filtering strategy used here parallels that used in the macro analysis. That is, the filter is applied so that it estimates the difference between the TV-NAIRU and the univariate trend in unemployment; this univariate trend is then added back in to obtain an estimate of the national TV-NAIRU. Specifically, as noted following equation 1.16, $\delta_t = \zeta_t - \beta u_t^N$. To maintain consistency with the treatment of the NAIRU in the macro regressions, rewrite this as $\delta_t + \beta u_t^* = \zeta_t - \beta(u_t^N - u_t^*)$, where u_t^* is the univariate trend in unemployment. Thus,

$$\delta_t + \beta u_t^* = \mu_t + \zeta_t, \tag{1.17}$$

where $\mu_t = -\beta(u_t^N - u_t^*)$.

Equation 1.17 has the same form as equation 1.9, in the sense that the intercept drift term μ_t arises from the difference between the NAIRU and the univariate trend in unemployment, except that equation 1.17 has no regressors (the observable and unobservable macro shocks are combined and contained in ζ_t). Accordingly, the national TV-NAIRU is estimated as $u_{t|T}^N = u_t^* + \mu_{t|T}/\hat{\beta}$ (see equation 1.12), where $\hat{\beta}$ is the estimate of β from the state regressions, $\mu_{t|T}$ is the Kalman smoother estimate of μ_t, and μ_t is modeled as following a random walk (as in equation 1.10), ζ_t is modeled as serially uncorrelated, and the dependent variable in equation 1.17 is $\hat{\delta}_t + \hat{\beta} u_t^*$, where $\hat{\delta}_t$ are the estimated time effects.

Instrumental-Variables Estimation Strategy

As is discussed in the next section, the state data on unemployment are obtained from the merged outgoing rotation groups (MORGs) of the CPS. Because many states have a small number of CPS respondents, these estimates are quite noisy, which leads to errors-in-variables bias. To avoid this bias, we use an IV approach. The MORG sample can be split into two independent samples, depending on whether the month is odd or even (although households appear twice in the MORG, the odd and even months have no households in common). Estimates from both the odd- and the even-month samples will be measured with error, but, because the samples are randomly drawn, the estimation error is independent in these two samples. Thus, one set of estimates can be used as an instrument for the other set of estimates. In particular, we use unemployment rates estimated from the even months as an instrument for unemployment rates estimated in the odd months, and vice versa. In some of our specifications (for example, those with lags of the dependent variable), the measurement error will be correlated between the independent and the dependent variables as well. Therefore, we replace all variables in the equation (both dependent and independent) with estimates from odd months and instrument with corresponding estimates from even months.[6]

Weighting the Observations

There is some ambiguity about whether the state regressions are best estimated by weighting the observations. The sampling error in the dependent variable will be smaller for larger states, but it is only one component of the error term, so the actual form of heteroskedasticity is unknown. Simple IV (two-stage least squares) has the virtue of taking no stand on the form of this heteroskedasticity and treats each state as an independent, equally useful experiment. On the other hand, weighting the observations can provide an approximate adjustment for this heteroskedasticity and, if implemented using employment weights, also produces estimates more directly related to aggregate coefficients; in particular, the aggregate NAIRU estimates constructed from the state data will reflect population weights. Because of this ambiguity, we report results using both weighted and unweighted observations, where the weights are given the values of state employment.

THE STATE DATA SET

Our state-level analysis relies on a data set containing annual observations on each of forty-eight states (excluding Alaska and Hawaii as well as the District of

Columbia) from 1979 to 1999. The annual data on each state were derived from a variety of sources, described in this section.

Data Derived from the Current Population Survey

We derive most of our variables, including annual estimates of wages, unemployment, and labor-force characteristics for each state, from the CPS MORGs. The MORG data are available from 1979 through 1999. Each month, one-quarter of the CPS sample (the outgoing rotation groups) is asked a variety of labor-force questions, for a total sample of over 300,000 individuals each year. For each individual who reports being in the labor force, the survey provides labor-force status (unemployed or not in the reference week), gender, race (white, black, other), marital status (married or not), and age. Education is reported in each year, but, because the format of the question changed in 1992, we have recoded the education variable into a set of ten consistent categories. Most recent industry of employment is reported by all individuals who have worked in the last five years, and we collapsed this information into eleven major industries.

For individuals who are currently working, we calculated hourly wage as usual weekly earnings divided by usual weekly hours. Earnings at the top code were multiplied by 1.5, and wages below the 1st percentile were set equal to the wage at the 1st percentile. We also calculated whether these persons were self-employed, were union members or covered by a union contract (available only since 1983), or worked as temporary help. To be considered a temporary-help worker, one had to report working in the personnel-supply-services industry and being paid by the hour. This definition is the same as that used by David Autor (forthcoming) and Lewis Segal and Daniel Sullivan (1997), but it is believed that at least 50 percent of temporary workers misreport their industry in the CPS. Finally, we calculated potential experience as age minus years of education minus 6, where, after 1991, years of education were imputed on the basis of the respondent's reported education category, race, and gender.

Using the individual-level data from the MORGs for all individuals who were in the labor force, we constructed state-level estimates for each year that were based on three samples: the full MORG sample, the respondents from even-numbered months, and the respondents from odd-numbered months. Households that appear in the MORG sample in even-numbered months do not appear in the odd-numbered months, so, as already noted, the estimation error in these two samples will be independent. In each sample, we constructed weighted estimates for each state and year (using weights provided by the CPS) of the unemployment rate and the fraction of the labor force in each age, education, race, and gender category. In addition, for employed individuals, we calculated the fraction of the workforce in each major industry, working in the temporary-help industry, self-employed, and covered by a union contract (or a union member). Finally, we calculated average and median log hourly wages for all workers, for hourly workers, and for full-time workers.

To construct state-level estimates of wages and unemployment that adjusted for changes over time in characteristics of the workforce, we estimated separate cross-sectional regressions for each year. In particular, for each year, we estimated a regression of either unemployment status or the log hourly wage on state fixed effects and controlled for ten education categories, three race categories, a quartic in experience, and an interaction between gender and all other regressors. In addition, controls for eleven major industries were included in the wage equation (but not in the unemployment equation). On the basis of this regression, each state's adjusted mean wage was predicted using that state's intercept and the average value of the covariates in the United States over the period 1979 to 1998 (calculated from the MORG sample).

Supplemental State-Level Data

In addition to the MORG data, we use a variety of labor-market measures that are available by state and year over most of our time period.[7] Data for each state on the unemployment-insurance replacement ratio are available through 1998 from the Information Technology Support Center unemployment-insurance website *(www.itsc.state.md.us)*. Data on the minimum wage have been derived from various issues of the *Monthly Labor Relations Review*. Data on the proportion of employment in the temporary-services sector come from county business patterns (these data are available only through 1996, so our specifications use them with a one-year lag to avoid losing observations). Finally, the proportion of the population age twenty-five to sixty-four on disability insurance and supplemental security income has been estimated from administrative data and provided to us by David Autor and Mark Duggan.

The Bureau of Economic Analysis website *(www.bea.doc)* was used to obtain data (available by state and major industry from 1978 to 1998) on gross state product (GSP) and total full-time and part-time employment (from table SA25), from which estimates of labor-productivity growth were derived. For each state, we constructed estimates of labor-productivity growth in two ways. Our primary method uses state-level estimates of GSP and employment and calculates labor-productivity growth in each year as $100[\ln(GSP_t/GSP_{t-1}) - \ln(employment_t/employment_{t-1})]$. Our secondary method is to estimate productivity growth in each state as a weighted average of the national-level estimates of labor-productivity growth in eleven major industries. National-level estimates of labor-productivity growth in each industry were derived from national estimates of GSP and employment as previously outlined. Each industry's productivity growth was weighted by the employment share in that industry (as estimated from the MORG data) in a given state and year to derive state-level estimates of labor-productivity growth.

EMPIRICAL STATE WAGE PHILLIPS CURVES

Benchmark Estimated Phillips Curves

Benchmark wage Phillips curve regressions of the form 1.16 are reported in table 1.6. These benchmark regressions do not include any structural variables (X_{it}) that might explain the movements in the NAIRU; these variables will be added in the eighth section.

Four features of table 1.6 are noteworthy. First, including state effects substantially changes the value of the coefficient on the unemployment rate. Although the state effects are usually jointly insignificant at the 5 percent level in these and subsequent regressions, excluding them evidently introduces omitted-variables bias into the estimated slope, so, henceforth, the state effects will be retained.

Second, using IV estimation to mitigate errors-in-variables bias leads to coefficients on the unemployment rate that are approximately one-third larger than the OLS estimates. This is consistent with the implication of the standard measurement-error model that the OLS estimator is biased toward 0. An important issue in IV regression is whether the instruments are correlated with the variable that they are instrumenting, that is, whether they are "weak." When there is a single variable being instrumented, this can be checked by seeing whether the F-statistic on the instruments in the first-stage regression is at least 10 (Staiger and Stock 1997). For the regressions in table 1.6, this first-stage F-statistic is always at least 100, which gives us confidence in applying standard asymptotic distribution theory to these IV regressions.

Third, the time effects are jointly significant in all specifications.

Fourth, the estimated slope of the state Phillips curve is large and statistically significant. In our preferred specification (IV with state and time effects), the estimated slope is -0.59, with a standard error of 0.11. This estimate is twice what we obtained using the quarterly macro data (-0.28 in column 1 of table 1.2), although it is comparable to the macro estimates in Gordon (1998). Because the state regressions control for both time and state effects, these results suggest

TABLE 1.6 / Phillips Curve Estimates from State Data
$\omega_{it+1} - \theta_{it+1} - \pi_{t+1} = \alpha_i + \delta_t + \beta u_{it} + \nu_{it+1}$

	OLS		IV	
Phillips curve slope (SE)	−.165	−.409	−.208	−.586
	(.045)	(.070)	(.061)	(.113)
Year effects p-value	.00	.00	.00	.00
State effects p-value	Excluded	.81	Excluded	.94
R^2	.26	.29	N.A.	N.A.

Notes: $N = 912$ (1979 to 1997). IV estimates use even months of MORG data as the instrument for odd months. N.A. = not applicable.

that idiosyncratic movements in a state's unemployment rate portend large idiosyncratic changes in the rate of growth of real wages.

Sensitivity Analysis

Tables 1.7 and 1.8 summarize results from several modifications to the benchmark specification. Table 1.7 shows results from changes in the estimation procedure (weighting the observations and alternative IV estimators), changes in the wage and unemployment variables (adjustments for changing demographics, industry mix, and so on), changes in the measurement and treatment of productivity, and changes in the assumed timing of the variables. Table 1.8 summarizes results from specifications with more general dynamics. We discuss each of these in turn.

The first two rows of table 1.7 consider different estimators. The benchmark specification (IV with time and state effects, shown in the last column of table 1.6 and repeated at the top of table 1.7) used unweighted observations, with observations for even months of the MORG sample used as instruments for observations corresponding to odd months. Specification 1 in table 1.7 reverses this and uses odd months as instruments for the even months. Specification 2 weights the state obervations by state employment. In both specifications, there is little change from the benchmark model.

Specifications 3 to 9 use different measures of wages and the unemployment rate. Specifications 3 to 6 use adjusted wage and employment data. The adjusted measures were computed by controlling for education, experience, race, gender, and industry (depending on the variable and the specification), as detailed in the table notes, using the adjustment method described in the sixth section. The benchmark specification uses the sample average wages of all workers (as described in the sixth section), and specifications 7 to 9 consider alternative wage series (median wage of all workers, mean wage of full-time workers, mean wage of hourly employees). The slope coefficient on the unemployment rate is essentially unaffected by these modifications to the benchmark specification.

Specifications 10 to 12 modify the way in which productivity enters the model. Specification 10 omits productivity from the analysis, with the result that the dependent variable in the regression is the log of real-wage growth. Specification 11 adds productivity as a regressor, relaxing the constraint of a unit elasticity implicit in the benchmark model. Both these changes to the benchmark specification lead to slightly larger estimates of the slope coefficient. The estimated productivity elasticity from specification 11 (not shown) is small (0.17, with a standard error of 0.13). There are several possible explanations for this. This result is consistent with substantial measurement error in the state productivity data. Additionally, even state productivity, precisely measured, contains short-term fluctuations, and, if it is trend productivity rather than annual productivity that is reflected in wages (as we have assumed in the macro specifications), then these estimates would be further biased toward 0. Finally, it might

TABLE 1.7 / Alternative Phillips Curve Estimates
$\omega_{it+1} - \theta_{it+1} - \pi_{t+1} = \alpha_i + \delta_t + \beta u_{it} + \nu_{it+1}$

Deviation of Specification from Benchmark	PC Slope (SE)
(0) None	−.586 (.113)
(1) Odd month IV for even months	−.447 (.113)
(2) Weighting by size of state workforce	−.542 (.089)
(3) Adjusted Wages[2] and Unemployment Rate[1]	−.506 (.097)
(4) Adjusted Unemployment Rate[1]	−.535 (.111)
(5) Adjusted Unemployment Rate[2]	−.534 (.129)
(6) Adjusted Wages[2]	−.552 (.098)
(7) Wages–Median/All	−.625 (.151)
(8) Wages–Mean/Full-Time	−.526 (.128)
(9) Wages–Mean/Hourly	−.660 (.119)
(10) Productivity growth omitted	−.637 (.092)
(11) Productivity growth used as a regressor	−.654 (.094)
(12) Weighted industry productivity growth	−.718 (.095)
(13) Productivity growth dated t	−.418 (.114)
(14) Unemployment dated $t + 1$	−.385 (.112)

Notes: The table shows estimates of the Phillips curve slope and the associated standard errors for alternative specifications. The stub column describes the deviation from the benchmark model (given in the last column of table 1.6 and labeled [0] in this table). Specification (1) reverses the role of instruments and regressors in the baseline specification. Specification (2) uses weighted IV, where the weights are the size of the state workforce. Specifications (3) to (6) use adjusted values for wages and/or the unemployment rate. Variables with a "1" superscript were adjusted using cross-sectional regressions each year that controlled for ten education categories, three race categories, a quartic in potential experience, and an interaction between gender and all other regressors. Variables with a "2" superscript used these regressors together with eleven major-industry indicator variables. Specifications (7) to (9) use alternative wage measures: median hourly wage for all workers; mean hourly wage for full-time workers; and mean hourly wage for workers paid on an hourly basis. Specification (10) uses the dependent variables $\omega_{it+1} - \pi_{t+1}$, (11) adds θ_{it+1} as a regressor (relaxing the unit elasticity constraint), and (12) uses a weighted industry measure of state productivity growth. Specification (13) used the dependent variable $\omega_{it+1} - \theta_{it} - \pi_{t+1}$, and specification (14) uses u_{it+1} as the regressor. (The results are invariant to the timing of π_t because of the inclusion of time effects.) The sample period was 1979 to 1998 (N = 960) for specification 10 and 1979 to 1997 (N = 912) for all other specifications.

be that productivity affects wage growth only with a lag, an explanation that we investigate subsequently.

Specification 12 uses an alternative measure of state productivity, the weighted average industry productivity growth that was described in the sixth section. This change produces a further steepening of the estimated Phillips curve.

The final two specifications in table 1.7 change the dating of productivity growth and the unemployment rate. Specification 13 replaces θ_{it+1} with θ_{it}, and specification 14 replaces u_{it} with u_{it+1}. Both changes result in a somewhat smaller estimated slope, suggesting that the timing convention in the benchmark specification is appropriate.

CHANGES IN THE DYNAMIC SPECIFICATION Table 1.8 considers changes in the dynamic specification of the model. Panel A allows the lagged level of productivity-adjusted real wages to enter the regression, relaxing the unit root constraint implied by the Phillips curve specification. We show results for both OLS and IV, specifications that include and exclude the productivity adjustments to real wages, and specifications that include the demographic and industry-mix adjustments to wages. The OLS estimates show some evidence supporting mean reversion in real wages: the coefficients on lagged real wages range from -0.09 to -0.15 (implying AR roots of 0.85 to 0.91 for annual real wages) and appear to be statistically significant. However, measurement error in the MORG wage series implies that the OLS estimators have a negative bias. When IV is used to eliminate this bias, the estimated coefficients on lagged wages are much closer to 0 and are statistically insignificant in the specifications that include productivity adjustments. (For similar estimates, see Blanchard and Katz 1997.)

Panel B of table 1.8 summarizes results for two specifications that allow distributed lags of wage growth and the unemployment rate in the model. These changes have little effect on the estimated slope of the Phillips curve, and the additional lagged variables are jointly insignificant in the regression.

SUMMARY OF THE SENSITIVITY ANALYSIS The benchmark specification relates annual observations of productivity-adjusted real-wage growth to unemployment rates one year earlier together with time and state effects. The estimated slope is approximately -0.6, and there are no additional dynamics. The results presented in tables 1.7 and 1.8 are broadly consistent with this specification.

National TV-NAIRUs Estimated Using the State Data

Estimates of the national TV-NAIRU derived from the estimated time effects in the benchmark model are plotted in figure 1.8. Also shown are results from the specification that omits productivity and uses the real wage instead of unit labor cost as the dependent variables (specification 10 in table 1.7). These estimates were computed by the method described in the fifth section. The figure also plots the national unemployment rate and its univariate trend. The means of the

TABLE 1.8 / Alternative Specification of Phillips Curve Dynamics

A. Relaxing the Unit Root Constraint on Real Wages

$\omega_{it+1} - \theta_{it+1} - \pi_{t+1} = \alpha_i + \delta_t + \beta u_{it} + \gamma[\ln(W_{it}/P_t) - \ln(\text{productivity}_{it})] + \nu_{it+1}$

Variable	OLS			IV		
Unemployment rate	−.401	−.375	−.480	−.581	−.548	−.612
	(.065)	(.062)	(.049)	(.113)	(.097)	(.089)
Real wage	−.103	−.090	−.146	−.010	−.028	−.057
	(.021)	(.022)	(.015)	(.034)	(.030)	(.026)
Productivity adjustment	Yes	Yes	No	Yes	Yes	No
Adjusted Wages²	No	Yes	No	No	Yes	No

B. Allowing Distributed Lags

$\omega_{it+1} - \theta_{it+1} - \pi_{t+1} = \alpha_i + \delta_t + \beta u_{it} + \alpha_\omega(L)[\omega_{it} - \theta_{it} - \pi_t] + \alpha_u(L)\Delta u_t + \nu_{it+1}$

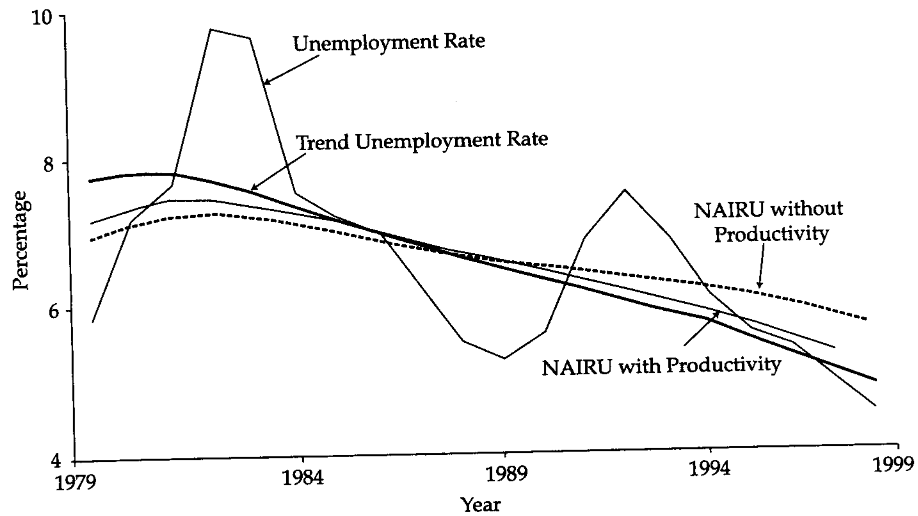
Variable	Coefficient (SE)	
u_{it}	−.486	−.549
	(.131)	(.161)
$\omega_{it} - \theta_{it} - \pi_t$	.051	.031
	(.103)	(.107)
$\omega_{it-1} - \theta_{it-1} - \pi_{t-1}$		−.210
		(.094)
Δu_t	−.404	−.141
	(.345)	(.411)
Δu_{t-1}		−.020
		(.520)
p-value for $\alpha_\omega(L) = 0$	.62	.08
p-value for $\alpha_u(L) = 0$	.24	.86

Notes: For panel A, specifications with "no" productivity adjustment omit the terms θ_{it+1} from the dependent variable and $\ln(\text{productivity}_{it})$ from the regressor. For the definition of Adjusted Wages², see the notes to table 1.7. For panel B, the sample period is 1981 to 1997 ($N = 816$).

dependent variables in the two state regressions differ, so, for comparison purposes, the state estimates of the national TV-NAIRUs have been shifted so that they have the same mean as the national unemployment rate over the period 1979 to 1997.

The estimated national TV-NAIRU based on the state unit labor-cost regression is similar to the univariate trend in national unemployment and, thus, is similar to the TV-NAIRUs estimated using the macro data. In contrast, the estimated national TV-NAIRU based on the state real-wage regression falls only slightly in the 1990s. Mechanically, this arises because the real-wage regressions implicitly introduce drift in the mean productivity-growth rate into the NAIRU: the sharp increase in real wages in the late 1990s implies that the unemployment

FIGURE 1.8 / State Estimates of the TV-NAIRU



Sources: Bureau of Labor Statistics and authors' calculations.

rate must have been well below the NAIRU during this period, if one neglects increases in the growth rate of productivity. On incorporating productivity into the dependent variable in the state panel regressions, the NAIRU falls substantially. This, then, is consistent with the conclusion from the analysis of the macro data that incorporating productivity reconciles the strong real-wage growth of the 1990s with an estimate of a declining unemployment TV-NAIRU estimated in price regressions.

We conclude from these estimates that, despite substantial differences in the data sets, span, and periodicity, the state evidence and the macro evidence are mutually consistent. This sanguine conclusion must be tempered by a recognition that both state estimates of the TV-NAIRUs are quite imprecise: the Kalman smoother estimate of the standard error on the decline in the state estimate of the TV-NAIRU from 1992 to 1997 using unit labor costs is 0.7 percentage points, and this standard error ignores estimation error. This imprecision should not be too surprising because the national estimates are based on smoothing the time series of estimated time effects, which has only twenty annual observations.

STATE EVIDENCE ON STRUCTURAL SOURCES OF SHIFTS IN THE NAIRU

This section reports the results of state panel regressions that examine the stability of the Phillips curve over time and across regions and that include variables that represent possible structural factors that determine the NAIRU. Our conclusions from the state regressions are summarized at the end of this section.

Stability

The stability of the wage Phillips curves over time and across regions is investigated in table 1.9. There is no evidence that the slope of the Phillips curve has changed over time. This is shown in table 1.9 using productivity-adjusted real-wage growth; similar results, not reported in the table, were obtained using real-wage growth without a productivity adjustment. This confirms Katz and Krueger's (1999) finding (obtained using OLS) that the state Phillips curve has been stable over time.

Interestingly, there is some evidence that the coefficient on the unemployment rate differs depending on region of the country (where regions are defined to be the four census regions). The Northeast and the West are estimated to have flatter Phillips curves than other regions do. These regional interactions are marginally significant (the p-value is .07). Understanding whether there actually is regional variation in this slope, and, if so, why, is an interesting topic for future research.

TABLE 1.9 / Stability of the Phillips Curve Through Time and Space
$\omega_{it+1} - \theta_{it+1} - \pi_{t+1} = \alpha_i + \delta_t + \beta_{it}u_{it} + v_{it+1}$

Variable	Coefficient (SE)		
u_{it}	−.586 (.113)	−.623 (.142)	−.736 (.139)
$u_{it} \times 1(t \leq 1984)$		.066 (.143)	
$u_{it} \times 1(t \geq 1992)$		.014 (.204)	
$u_{it} \times 1(i$ in Northeast)			.509 (.198)
$u_{it} \times 1(i$ in North Central)			.065 (.159)
$u_{it} \times 1(i$ in West)			.214 (.203)
Temporal stability p-value		.895	
Spatial stability p-value			.065

Notes: The sample period is 1979 to 1997 (N = 912). The temporal stability p-value is associated with the Wald test for the hypothesis that the coefficients on u_{it} interacted with the time indicators are zero. The spatial stability p-value is associated with the Wald test for the hypothesis that the coefficients on u_{it} interacted with the region indicators are 0.

Demographics and Education

We now turn to regressions that investigate possible structural reasons that the NAIRU might change over time. The first such regressions examine the role of demographics and education, a theme recently emphasized by Robert Shimer (1998).

Table 1.10 reports the results of regressions that include, either individually or together, the percentage of high school dropouts, the percentage of college graduates, the percentage white in the work force, the percentage female in the workforce, and the percentage of the workforce between twenty-five and fifty-four years of age. These variables and the unemployment rate are measured in percentage points. The estimated coefficients are large, but so are their standard errors, and none of the demographic or education variables are statistically significant at the 5 percent level, either individually or jointly. For example, the coefficient on the percentage of high school dropouts in the second column is −0.099, which implies that a 1 percentage point increase in the fraction of high school dropouts is associated with a decrease in the NAIRU of 0.15 percentage points (−0.099/0.642). However, the standard error of this estimated effect is very large (0.23 percentage points), with the result that this specification produces a 95 percent confidence interval for the effect of a 1 percentage point increase in high school dropouts on the NAIRU that ranges from a decline of 0.6

TABLE 1.10 / Demographic Variables and the Phillips Curve

$$\omega_{it+1} - \theta_{it+1} - \pi_{t+1} = \alpha_i + \delta_t + \beta u_{it} + \gamma X_{it} + \nu_{it+1}$$

Variable	Coefficient (SE)					
Unemployment rate	−.586	−.642	−.550	−.604	−.542	−.526
	(.113)	(.147)	(.134)	(.120)	(.159)	(.320)
Percent high school dropout		−.099				−.081
		(.140)				(.189)
Percent college graduate		.299				.577
		(.345)				(.918)
Percent white			.233			.164
			(.543)			(2.000)
Percent female				−.328		.030
				(.621)		(1.378)
Percent age twenty-five to fifty-four					−.113	−.381
					(.285)	(.981)
Education variables *p*-value		.339				.704
All demographics *p*-value						.712

Note: The sample period is 1979 to 1997 ($N = 912$).

percentage points to an increase of 0.3 percentage points. The large standard errors reflect the fact that, once state and time effects have been removed, there is only limited within-state variation in these slowly moving demographic variables.

Robert Lerman and Stefanie Schmidt (1999) have claimed that the recent rise (and impending decline) in the share of prime-age workers and the continuing rise in the share of college-educated workers are key factors in understanding recent (and future) changes in the labor market.[8] However, the timing of these demographic shifts does not coincide well with the downturn in the estimated NAIRU since 1992. In particular, the share of the workforce with a college degree increased steadily in both the 1980s and the 1990s, which would suggest a steady decline in the NAIRU over the entire period. Similarly, the share of the workforce between the ages of twenty-five and fifty-four increased dramatically between 1979 and 1992 but has been flat or has even declined since then. Thus, neither of these factors increased at a higher rate in the 1990s, which is what would be needed to explain the sudden decline in the NAIRU after 1992. Despite the inability of these demographic shifts to explain recent changes in the NAIRU, it is worth noting that the point estimates imply that the impending decline in the share of prime-age workers over the next twenty years (as the baby boomers retire) would exert upward pressure on the NAIRU. Thus, reconciling the state-level estimates of demographic effects on the NAIRU with the macro-level evidence of changes in the NAIRU is an important topic for future research.

Industry Characteristics and Temporary Help

Table 1.11 reports results for regressions including industry characteristics and the relative size of the temporary-help industry in the state. The results suggest that increases in the share of retail trade and services are associated with increases in the NAIRU, relative to manufacturing, but these effects are not statistically significant. Similarly, the point estimates suggest that increases in temporary help and self-employment lead to declines in the NAIRU, but, again, these estimated effects are not statistically significant.[9]

Government Policies

Table 1.12 examines the effect of state labor-market policies on the NAIRU. These policies include the percent of the prime-age workforce on disability insurance and supplemental security income, the minimum wage, the growth of the minimum wage, and the unemployment insurance replacement ratio. None of the coefficients on these regressors are statistically significant at the 5 percent level in any of the specifications.

TABLE 1.11 / Industry Characteristics and the Phillips Curve
$\omega_{it+1} - \theta_{it+1} - \pi_{t+1} = \alpha_i + \delta_t + \beta u_{it} + \gamma X_{it} + \nu_{it+1}$

Variable	Coefficient (SE)					
Unemployment rate	−.586	−.823	−1.084	−.599	−.575	−.854
	(.113)	(.170)	(.397)	(.115)	(.115)	(.179)
Percent durable mfg.		.333				.196
		(.227)				(.283)
Percent nondurable mfg.		−.166				−.247
		(.303)				(.312)
Percent retail trade		.275				.161
		(.708)				(.715)
Percent services		1.045				1.113
		(.501)				(.526)
Percent temp. help (CPS)			−19.239			
			(13.567)			
Percent temp. help (CBP)				−.290		−.159
				(.324)		(.383)
Percent self-employed					−.196	−.437
					(.272)	(.487)
Major industry p-value		.153				.164
All industry variables p-value						.312

Note: The sample period is 1979 to 1997 (N = 912).

TABLE 1.12 / Government Policy Variables and the Phillips Curve
$\omega_{it+1} - \theta_{it+1} - \pi_{t+1} = \alpha_i + \delta_t + \beta u_{it} + \gamma X_{it} + \nu_{it+1}$

Variable	Coefficient (SE)				
Unemployment rate	−.586	−.576	−.603	−.591	−.600
	(.113)	(.120)	(.114)	(.114)	(.122)
Percent of twenty-five to sixty-four on DI		.360			.376
		(.606)			(.611)
Percent of twenty-five to sixty-four on SSI		.046			.025
		(.577)			(.578)
Minimum wage (1999 dollars)			−.414		−.354
			(.864)		(.879)
Percent growth in minimum wage			−.085		−.084
			(.049)		(.049)
UI replacement rate				.015	.017
				(.043)	(.043)
DI/SSI variables p-value		.746			.749
Min. wage variables p-value			.207		.210
All policy variables p-value					.566

Notes: The sample period is 1979 to 1997 (N = 912). DI = disability insurance. SSI = supplemental security income. UI = unemployment insurance.

Summary of State Results

The finding that the unemployment rate significantly enters equations for either the change in labor share or change in real-wage equations is highly robust to changes in specification, including using different lags, using different measures of wages, using demographically adjusted data, and controlling for a large number of possible structural determinants of the NAIRU. The coefficient is estimated to be approximately -0.6. This coefficient is stable over time, although there seem to be some intriguing but uninvestigated variations in this coefficient over different regions of the country.

In contrast, the state evidence on the determinants of the NAIRU is generally negative. None of the state labor-market-policy variables are statistically significant. Neither industry composition nor the size of the temporary-help sector is a statistically significant determinant of the NAIRU in the IV regressions. Of the demographic and education variables, there is some evidence that the presence of relatively more workers in the prime-age work group contributes to reductions in the NAIRU, but this effect is estimated very imprecisely.

Despite the lack of statistical significance of these effects, it is worth checking whether the point estimates of the coefficients are consistent with these variables potentially having an economically substantial effect on the NAIRU and, in particular, on the evolution of the NAIRU through the 1990s. The point estimates associated with the demographic variables are, in fact, consistent with these variables having economically large effects, but the timing of these effects is inconsistent with their playing an important role in the fall in the NAIRU since 1992; rather, they contribute to a decline in the NAIRU prior to 1992, but their contribution is estimated to have increased the national NAIRU since 1992.

The point estimates associated with the industry-mix variables, including temporary help (the last column of table 1.11), do point to a contribution that would lower the NAIRU by approximately 0.5 percentage points from 1992 to 1997, but this effect is not statistically significant.

Overall, these state panel regressions fail to pinpoint any economic determinants of the TV-NAIRU. This accords with the macro evidence presented in Stock (forthcoming) that education and demographic variables are inconsistent at the macro level with the trends in the NAIRU. Our finding contrasts somewhat with that of Katz and Krueger (1999), who find some evidence that the rise of the temporary-help industry has contributed to the fall in the national NAIRU (they estimate that temporary help has reduced the national TV-NAIRU by approximately 0.4 percentage points since 1990). Our specifications differ somewhat from theirs, however, and we find that the effect of temporary help is not robust.

CONCLUSIONS

We assess these results by returning to the three questions of the introduction. First, did the Phillips curve break down in the 1990s, or did it simply shift, with a new and evolving NAIRU? In both our macro and our state analyses, we found abundant evidence of a stable relation both between the change in price inflation and economic activity gaps and between unit labor costs and these activity gaps. These two pieces of evidence are confirmatory and are particularly striking because the state panel estimates included time effects, thereby eliminating the variation in the data that drives the macro estimates. Moreover, in the macro analysis, we found evidence that there was little intercept drift when the regressions were estimated as gaps; that is, when specified in terms of gaps, neither the price nor the wage Phillips curve has shifted in the 1990s. Thus, this evidence falls squarely on the side of "the Phillips curve is alive and well but . . ." theories.

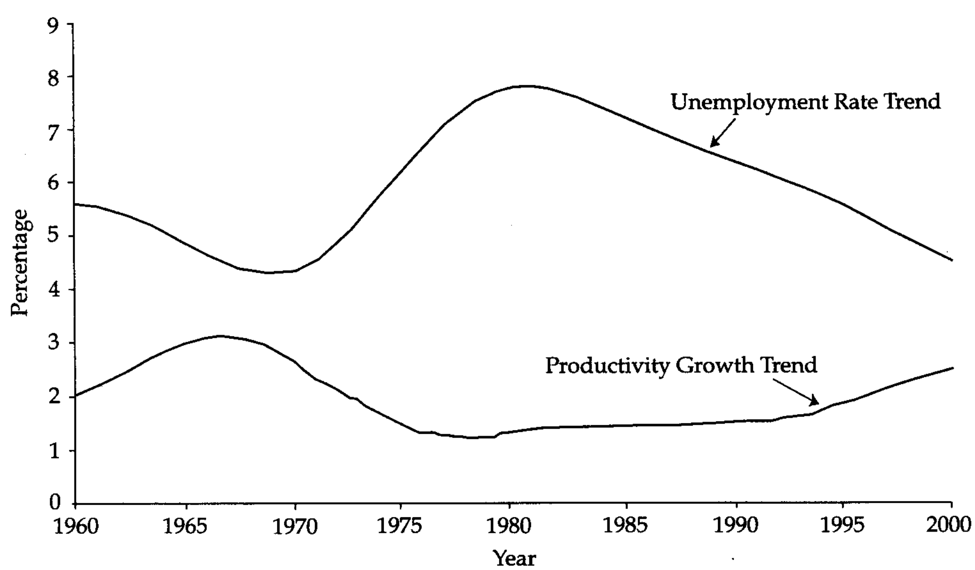
Second, given this finding, why has the price Phillips curve shifted in? The macro analysis suggests that the answer is *not* that we have had a particularly fortunate string of supply shocks; we did in the mid-1990s, but they were subsequently reversed. In addition, both the macro evidence and the state evidence suggest that an autonomous shift in wage- or price-setting behavior by firms is also not an important part of the inflation-unemployment story of the 1990s: had changes in price- or wage-setting behavior been the reason for the apparent decline in the NAIRU, the NAIRU and the univariate trend in unemployment would differ, but they do not. This finding is reinforced by the limited role that we found for industry-mix variables in the state-panel-data analysis. This suggests that labor-market factors, such as demographics, the rise of the temporary-help industry, or labor-market policies, must be the source of the changes in the NAIRU. Curiously, however, our results do not point in this direction. The macro estimates of the change in the NAIRU from 1992 to 2000 derived using the demographically adjusted unemployment rate are virtually the same as those derived using the total unemployment rate. Similarly, our attempts to identify structural determinants of the NAIRU using the state data were disappointing. Several effects pointed in the right direction, but the state data did not provide precise estimates of these effects.

Third, can the labor-productivity gains of the 1990s explain the apparently aberrant recent behavior of real wages as displayed in figure 1.2? Yes. The TV-NAIRUs estimated both on the macro and on the state data are the same whether the change in inflation or the changes in real wages less productivity is the dependent variable and differ sharply during the 1990s only when the productivity component is omitted from wages.

The simplest summary of these results is that, once one accounts for the univariate trends in the unemployment rate and in productivity, the 1990s present no price or wage puzzles. Thus, the task is to explain trend movements in productivity and in unemployment. In our framework, trend unemployment and the NAIRU are identified separately, but, as it happens, these two series track each other very closely. Because the long-run trend components of the rates of inflation and unemployment are essentially unrelated in the postwar U.S. data (Stock and Watson 1999a), we are skeptical of explanations that link the two, such as those of Akerlof, Dickens, and Perry (1996) or Taylor (2000).

It is potentially more promising to consider explanations that directly link the trend components of productivity growth and the unemployment rate. These univariate trends, which are plotted in figure 1.9, show a striking and intriguing negative correlation. The recent coincident increase in productivity growth and decrease in the unemployment rate recall a similar pattern in the early 1960s. This pattern reversed itself in the 1970s, when trend productivity growth fell and the trend unemployment rate increased. The close relation between the series can also be seen in the scatterplot of the series shown in figure 1.10. Of course, the figure must be viewed with caution since unrelated trends can spuriously appear to be correlated. Indeed, it seems reasonable to think of figure 1.10 as four data observations—1960 to 1967, 1967 to 1980, 1980 to 1993, and 1993 to 2000—and it is difficult to be sure of a correlation with only four observations. Yet we find the empirical results strong enough, and the question important enough, to warrant further attention by both macro and labor economists.

FIGURE 1.9 / Trend Unemployment and Productivity Growth



Source: Authors' calculations.

FIGURE 1.10 / Unemployment Rate and Productivity-Growth Trends

Source: Authors' calculations.

DATA APPENDIX

This appendix documents some features of the macroeconomic data used in the first four sections of this paper. All series are from the DRI Basic Economics Database (formerly Citibase). Quarterly averages were used for variables that were available monthly. Table 1A.1 lists the database mnemonic, a brief series description, and the series abbreviation used in tables 1.2 to 1.4.

We thank David Autor for helpful discussions and William Dickens and Robert Solow for detailed comments on an earlier version of this paper.

TABLE 1A.1 / Definitions of the Macroeconomic Data

Price series	
GDPD	Gross domestic product: implicit price deflator (GDP Def.).
GMDC	PCE, implicit price deflator (PCE Def.).
PUNEW	CPI-U all items CPI (CPI).
LBGDPU	Nonfarm business: implicit price deflator (NFB Def.).
Wage series	
LBPUR	Compensation per hour, employees: nonfarm business (Comp./Hr).
LEH	Average hourly earnings of production workers: total private nonagricultural (AHE PW).
LCP	Employment cost index (compensation): private industrial weeks (ECC-C).
LWI	Employment cost index (wage and salary): private industrial weeks (ECC-WS).
Real activity variables	
LHUR	Unemployment rate: all workers, sixteen years and over (Civ. Unemp.).
HSBR	Housing authorized: total new private housing units (Bldg. Perm.).
IPXMCA	Capacity-utilization rate: manufacturing (Cap. Util.).
Productivity	
LBOUTU	Output per hour of all persons: nonfarm business.
LOUTM	Output per hour of all persons, index—manufacturing.
Other variables	
PUXX	CPI-U: all items less food and energy.
EXVUS	Foreign exchange value of the U.S. dollar.
Constructed variables	
Exchange rate	EXVUS from 1973:1 through 2000:6. From 1959:1 to 1972:12 this is a trade-weighted average of the dollar exchange rates for France, Germany, Italy, Japan, and the United Kingdom, described in Stock and Watson (1989). The two series were linked in 1973:1.
Relative price of food and energy	$\ln(\text{punen}w_t/\text{pune}w_{t-1}) - \ln(\text{puxx}_t/\text{puxx}_{t-1})$.

TABLE 1A.1 / *Continued*

Wage- and price-control variable	This variable takes on the values .8 from 1971:3 to 1972:3, −.4 from 1974:2 to 1974:2, −1.6 from 1974:3 to 1974:4, −.4 from 1975:1 to 1975:1, and 0 for all other dates. It is taken from Gordon (1982).
Supply shocks	The specifications that included supply shocks included four lags of the relative price of food and energy, four lags of the log difference of the exchange rate, and the wage- and price-control variable.
Demographically adjusted unemployment rate	See the text.
Bias corrections for the inflation series	(CPI) inflation was adjusted for the improvements in measurement implemented by the Bureau of Labor Statistics, as suggested in Gordon (1998). The adjustments were taken from Council of Economic Advisers (1998, table 2.4, p. 80). The values used are −.12 from 1995:1 to 1995:4, −.22 from 1996:1 to 1996:4, −.28 from 1997:1 to 1997:4, −.49 from 1998:1 to 1998:4, and −.69 from 1999:1 to 2000:2.

Source: Authors' compilation.

NOTES

1. Application of the Horvath-Watson (1995) test rejects the null hypothesis of the noncointegration of these two series in favor of the alternative, that they are cointegrated with the cointegrating vector of (1, −1) at the 1 percent significance level (the test statistic was computed with four lags).
2. Gordon's (1998) derivation differs from ours, and he does not discuss cointegration explicitly. However, by imposing sums of coefficients in his lag polynomials to equal 1 (which he does in his empirical work), the resulting system is equivalent to equations 1.2 and 1.4, which in turn implies that the system is cointegrated.
3. Garner (1994) pointed out the stability of Phillips curves specified with capacity utilization and estimated with data through the early 1990s. Gordon (1998) and Stock (1998) estimated TV-NAIRUs for capacity utilization and found that they were quite stable compared with TV-NAIRUs for the unemployment rate. Our results confirm these findings and extend them through the end of the 1990s.
4. W. C. Brainard and George Perry (2000) estimate wage and price Phillips curves allowing for time variation in the coefficients and also find little change in the estimated slope. Their specification differs from ours in several respects; most notably, their equations are estimated using the levels of price and wage inflation, while we use the

change in price inflation and the change in productivity-adjusted real wage. Although their estimates suggest that the coefficients on lagged inflation have changed over time, they do not provide standard errors or any evidence of whether these changes are statistically significant. Doing so in their specification would require handling the persistence of their regressors in addition to the usual issues arising in time-varying coefficients.

5. The difference between the levels of the NAIRUs in price and wage Phillips curves reflects the decline in labor's share over the sample period computed using nonfarm-business wages, nonfarm-business productivity, and the GDP price deflator. The magnitude of this decline is evident from the tenth row of table 1.1. However, as the last row of table 1.1 shows, there is a much smaller decline when the nonfarm-business price deflator is used. Consequently, the levels of the NAIRU for the wage Phillips curves obtained using the nonfarm-business deflator are closer to those for the price Phillips curves.
6. This instrumenting strategy is simple but statistically inefficient. For an alternative, more efficient method, see Autor and Staiger (2001).
7. We thank David Autor for providing us with many of these data.
8. Lerman and Schmidt (1999) also present some limited evidence suggesting that there is no relation between the unemployment rate and wage growth in the late 1990s, which is at odds with the evidence that we present in table 1.6. In particular, they report, on the basis of 3 months of CPS data from 1995 and 1998, no relation between growth in wages at the state level from 1995:1 to 1998:1 and a state's unemployment-rate quartile in 1998:1. A number of aspects of the Lerman and Schmidt evidence are likely to bias their estimate toward finding no relation. In particular, the small sample sizes that result from using only three months of data exacerbate the measurement-error issues in the unemployment rate. More important, their focus on a single three-year difference with unemployment measured at the end of the difference is at odds with the usual Phillips curve specification, which focuses on short differences and lagged unemployment.
9. Katz and Krueger (1999) find a statistically significant effect of temporary help on the NAIRU in Phillips curves estimated from state data. Their result appears to depend in an important way on the particulars of their specification and estimator (OLS).

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