The Impact of Evolving Labor Practices and Demographics on U.S. Inflation and Unemployment

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Abstract

Since the early 1990s, NAIRU estimates have declined and unemployment duration has risen relative to the unemployment rate. These developments may have arisen from the aging of the workforce or practices reducing job turnover. We assess the internal consistency of these hypotheses using simulation methods and test their external consistency using modified NAIRU models. We find that demographics cannot fully account for changes in the NAIRU, consistent with Staiger, Stock, and Watson (2001) and in contrast to Shimer (1998, 2001). Instead, our results attribute shifts in the NAIRU and duration to a combination of shifts in demographics and job turnover.

JEL Codes: J30, J64, E31
Key Words: unemployment, duration, NAIRU

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I. Introduction

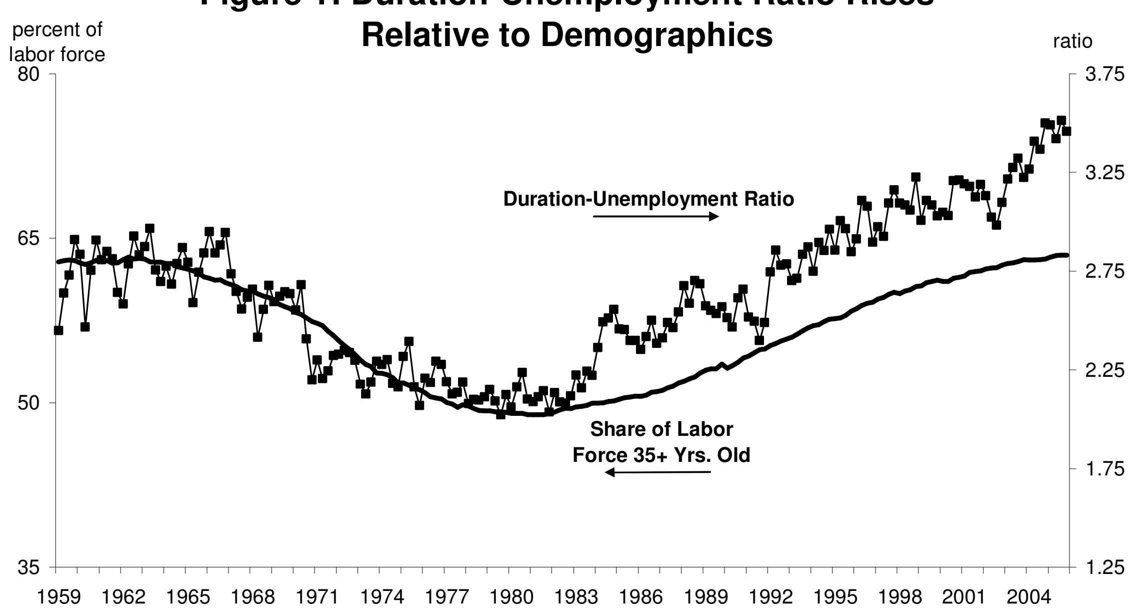
This study assesses whether an apparent decline in the NAIRU since the early 1990s is attributable to changes in U.S. labor practices and demographic factors that are reflected in a concomitant rise in average unemployment duration relative to the unemployment rate. In this period, the U.S. unemployment rate fluctuated in a low range, with inflation below what prior experience suggested, sparking a renewed recognition that the "natural" rate of unemployment is not precisely known and is subject to shifts. Although this phenomenon has been well documented (e.g., Staiger, Stock, and Watson, 1997, and Gordon, 1997), there has been less progress in accounting for the drop in the NAIRU. While demographic changes have played a role (Abraham and Shimer, 2001), they cannot fully account for the decline (Staiger, Stock, and Watson, 2001).

Demographics also cannot largely explain the rise in the ratio of the average duration of unemployment to the unemployment rate, which has increased more since the late 1990s than what prior experience would suggest based on swings in the share of the labor force under the age of 35 (Figure 1).

This study argues that both of these phenomena may have resulted from a fall in the rate of job turnover. A decrease in job turnover means that there are fewer job vacancies at a given unemployment rate, implying that the hiring rate is lower at each unemployment rate. Thus, a fall in the hiring rate results in an increase in the ratio of average duration to unemployment. In addition, if firms pay efficiency wages, a fall in the hiring rate at each unemployment rate implies that the profit-maximizing wage is lower, resulting in a lower equilibrium unemployment rate for the economy.

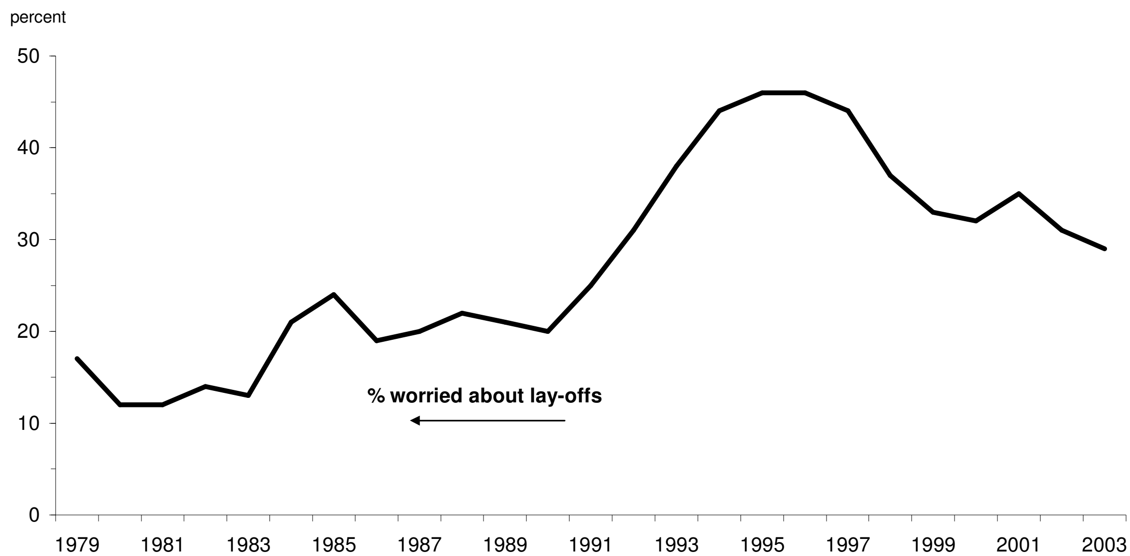
Shimer (2005) presents evidence that there has been a decrease in job separation since the early 1990's. While it is difficult to pinpoint the exact reason for a fall in job turnover, there are several plausible explanations for this decrease. One possible reason is that the downsizing that occurred during the 1990-91 recession and its aftermath caused workers to perceive a lower degree of job security. Workers who are more worried about layoffs may be less likely to change jobs, since they may be uncertain about the quality of the match with their new employer and since the first workers to be laid off are often the ones most recently hired. Consistent with this explanation, surveys of workers at large firms conducted by International Survey Research show that the share of workers worried about their job security has generally been higher since the late 1980s (Figure 2). A second possible reason for a fall in the separation rate is that women's labor force attachment has increased, as suggested by Abraham and Shimer (2001). A third explanation is that selection problems in obtaining health benefits at a new job have become more significant as real health care costs have risen. This, in turn, may have lowered turnover by discouraging workers from quitting and seeking new job matches.

Figure 1: Duration-Unemployment Ratio Rises Relative to Demographics



Sources: Bureau of Labor Statistics and authors' calculations.

Figure 2: Percent of Workers Frequently Worried About Being Laid Off



Source: International Survey Research .

We use two approaches to analyze the validity and plausibility of the hypothesis that the fall in the natural rate and the rise in the duration-unemployment ratio are related to a decline in the rate of job turnover. First, we assess whether changing labor practices and demographics can account for a falling NAIRU and rising duration-to-unemployment ratio using an efficiency wage framework. Second, we add the duration-to-unemployment rate ratio to NAIRU models to see if this labor gauge adds marginal information in the presence of the overall or demographically-adjusted unemployment rate. The rationale for including the duration-unemployment ratio is that changes in job turnover at a given unemployment rate would be reflected in movements in this ratio and changes in job turnover at each unemployment rate affect the NAIRU. Thus, the unemployment-duration ratio serves as a proxy for the job turnover rate (holding the unemployment rate constant) and should add marginal information in estimating the NAIRU.

A related motivation for including the duration-unemployment ratio in Phillips curve equations is that, in efficiency wage models, the inflation rate may be related to the probability of finding a job. While the unemployment rate affects the probability of finding a job, it is not the sole determinant. The duration-unemployment ratio adds marginal information about the probability of finding a job and thus may add marginal information in predicting the inflation rate. Adding the unemployment-duration rate to NAIRU models also allows us to estimate how the NAIRU has evolved. In essence, we test the internal consistency of our hypothesis by using simulation methods in an efficiency wage model, while also checking its external consistency using regression analysis.

Results from both approaches support the view that labor markets have been affected at a macro level by both a decreased likelihood of worker-initiated job turnover and a shift toward a more experienced labor force pool. To establish these findings, section II presents an efficiency wage model and discusses the results of simulating shifts in labor practices and labor force composition. As a check on external consistency and to assess using an additional measure of overall labor slack, the third section augments NAIRU models with data on demographics and the duration of unemployment. Our last section concludes by summarizing and interpreting our findings.

II. Simulations of the Natural Rate and Average Duration

<u>Model</u>

This subsection describes the model used to analyze changes over time in the natural rate of unemployment and the duration-unemployment ratio. It is assumed that there are two classes of workers, where type 1 workers are older (age 35 and over) and type 2 are younger. We denote the number of each type as LF_1 and LF_2 and employment of each type as $\overline{L}_1$ and $\overline{L}_2$. Note that we could distinguish among workers using other criteria, such as white- versus blue-collar occupations, or males versus females. Thus, the model has the flexibility to analyze different types of issues related to the natural rate and duration of unemployment.

We assume firms set efficiency wages, where the efficiency of type *i* workers is:

$$e^i = e^i\left(\frac{W_i}{\overline{W}_i}, h(u_i, q_i)\right), \tag{1}$$

where W_i is the wage the firm pays, $\overline{W}_i$ is the economywide average wage rate for type *i* workers, and *h* is the probability of an unemployed worker being hired in a given period. It is assumed that the probability of hire depends negatively on the unemployment rate (u_i) and positively on the rate of separation into unemployment (q_i), where the latter reflects that there are more job openings when the separation rate is higher.[1]

The probability of hire can be calculated as follows. The number of type *i* unemployed workers at the beginning of a period (NUB_i) equals the number unemployed in the prior period plus separations at the end of the prior period. Thus,

$$\begin{aligned} NUB_{it} &= u_{i,t-1}LF_{i,t-1} + q_i\overline{L}_{i,t-1} \\ &= u_{i,t-1}LF_{i,t-1} + q_i(1-u_{i,t-1})LF_{i,t-1}. \end{aligned}$$

It is assumed that outflows from the pool of the unemployed (OPU_i) equal separations in the previous period (since these workers need to be replaced). Accordingly, outflows are

$$OPU_{i,t} = \alpha s_i\overline{L}_{i,t-1} = \alpha s_i(1-u_{i,t-1})LF_{i,t-1},$$

where s_i is the proportion of workers who separate either into unemployment or out of the labor force,[2] and α is the proportion of vacancies that are filled by individuals in the pool of the unemployed (as opposed to individuals who are not in the labor force). If it is assumed that a fraction, β, of separations are into unemployment (as opposed to out of the labor force), then $q_i=\beta s_i$. This relationship allows *OPU* to be expressed as

$$OPU_{i,t} = (\alpha / \beta) q_i (1 - u_{i,t-1}) LF_{i,t-1}.$$

Thus, the probability of an unemployed worker of type *i* being hired in a given period is

$$h_i = \frac{(\alpha / \beta) q_i (1 - u_{i,t-1})}{q_i (1 - u_{i,t-1}) + u_{i,t-1}}. \tag{2}$$

A firm's output (Q) depends on the quantity of each type of labor employed, with the following production function:

$$Q = f\left(e^1\left(\frac{W_1}{\overline{\overline{W}}_1}, h(u_1, q_1) \right) L_1, e^2\left(\frac{W_2}{\overline{\overline{W}}_2}, h(u_2, q_2) \right) L_2 \right).$$

The profits of a typical firm are given by the equation,

$$\pi = Pf\left(e^1\left(\frac{W_1}{\overline{\overline{W}}_1}, h(u_1, q_1) \right) L_1, e^2\left(\frac{W_2}{\overline{\overline{W}}_2}, h(u_2, q_2) \right) L_2 \right) - W_1 L_1 - W_2 L_2. \tag{3}$$

Differentiating (3) with respect to the employment of type *i* workers and the wages of type *i* workers and setting the derivatives equal to 0 yields

$$\frac{d\pi}{dL_i} = 0 = Pf_i e^i\left(\frac{W_i}{\overline{\overline{W}}_i},\, h(u_i, q_i)\right) - W_i, \qquad \text{and} \tag{4}$$

$$\frac{d\pi}{dW_i} = 0 = Pf_i e_W^i\left(\frac{W_i}{\overline{\overline{W}}_i},\, h(u_i, q_i)\right)\frac{1}{\overline{\overline{W}}_i}L_i - L_i. \tag{5}$$

By substituting (5) into (4), we obtain the following equation, which is analogous to the Solow (1979) condition:

$$\frac{e_W^i\left(\frac{W_i}{\overline{\overline{W}}_i},\, h(u_i, q_i)\right)}{e^i\left(\frac{W_i}{\overline{\overline{W}}_i},\, h(u_i, q_i)\right)} = 1. \tag{6}$$

The effect of the separation rate on the natural rate of unemployment can be calculated by setting $W_i / \overline{W_i}$ equal to 1 (since $W_i = \overline{W_i}$ in a steady-state equilibrium), and totally differentiating (6). Thus,

$$e_{Wh}^i h_u du_i + e_{Wh}^i h_q dq_i = e_h^i h_u du_i + e_h^i h_q dq_i,$$

yielding the following relationship:

$$\frac{du_i}{dq_i} = -\frac{h_q}{h_u} > 0.$$

From equation (2), the values of h_q and h_u are

$$h_q = \frac{(\alpha/\beta)u(1-u)}{[q(1-u)+u]^2}, \quad \text{and} \quad h_u = \frac{-(\alpha/\beta)q}{[q(1-u)+u]^2},$$

so that

$$du_i = \frac{u_i(1-u_i)}{q_i} dq_i . \qquad (7)$$

The average duration of unemployment can be calculated as follows. Let Z_i denote the number of workers of type i. In each period, the number of type i workers who have undergone a separation and remain unemployed is

$$\sum_{k=1}^{\infty} Z_i (1-h_i)^k = Z_i \frac{1-h_i}{h_i} ,$$

and the total number of periods these workers have been unemployed is

$$\sum_{k=1}^{\infty} kZ_i (1-h_i)^k = \sum_{k=1}^{\infty} \sum_{m=k}^{\infty} Z_i (1-h_i)^m = Z_i \frac{1-h_i}{h_i^2} .$$

Thus, the average duration of unemployment for this group of workers is given by the equation:

$$D_i = \frac{1}{h_i} = \frac{q_i(1-u_i)+u_i}{(\alpha/\beta)q_i(1-u_i)} . \qquad (8)$$

The average duration of unemployment depends on both q and u. Differentiating (8) with respect to these variables yields

$$dD_i = \frac{q_i du_i - u_i(1-u_i)dq_i}{(\alpha/\beta)q_i^2(1-u_i)^2} . \qquad (9)$$

Thus, the average duration of unemployment depends positively on u and negatively on q. Note that a change in q directly affects D as expressed in (9), and it also indirectly affects D through its effect on u. To find the total effect of q on D, we substitute (7) into (9), which yields

$$dD_i = \frac{q_i \dfrac{u_i(1-u_i)}{q_i} dq_i - u_i(1-u_i)dq_i}{(\alpha/\beta)q_i^2(1-u_i)^2} = 0 .$$

Thus, the indirect effect of q on D exactly offsets the direct effect, so that the total effect is 0.

At the aggregate level, the average duration of unemployment is

$$D = p_1 D_1 + p_2 D_2 ,$$

where p_i is the share of workers of type i among the unemployed.

<u>Calibration</u>

The model was calibrated with U.S. data from 1960-1970, and simulations were run over 1960-2005. As previously discussed, it is assumed that type 1 workers are age 35 and over and that type 2 workers are younger than age 35. This section discusses how parameters for the simulations were determined.

In the initial calibration, it is assumed that 61.3% of workers are type 1 and 38.7% of workers are type 2, in line with the age composition of the workforce over 1960-70. The unemployment rate is calibrated to match the average value of the natural rate of unemployment estimated by the Congressional Budget Office (CBO) over 1960-1970. In this period, the actual unemployment rate averaged 3.3% for older workers and 7.2% for younger workers. However, the actual average unemployment rate was below the CBO's average NAIRU estimate. To make the simulated unemployment rate compatible with the

CBO's average NAIRU estimate, the unemployment rates for the separate age groups are set at 3.89% and 8.49% in calibrating the model.

It is assumed that q differs across the two types of workers, with this variable higher for younger workers than for older workers, in line with evidence from Hall (1982) and Blanchard and Diamond (1990). To determine baseline values of q for each group of workers, we make use of the facts that the ratio of duration to unemployment averaged 2.75 over the 1960-70 period and that the average unemployment duration was 1.5 times longer for older workers than for younger workers. These conditions result in values of q_1=0.00195 per week (which corresponds to a monthly separation rate of 0.00845) and q_2=0.00686 per week (which corresponds to a monthly separation rate of 0.0295).[3] These figures imply an overall average monthly separation rate of 0.0164.[4]

Simulation results

After determining baseline values for u and q for each group of workers from the initial calibration, simulations were run with annual data from 1960-2005. In these simulations, the proportion of workers in each age category was determined from the actual percentage in Bureau of Labor Statistics data. In addition, within each broad category, the values of q_1 and q_2 were allowed to vary over time, depending on the age and gender composition of each broad age group. Among workers 35 and older, data from Blanchard and Diamond (1990) indicate that separations are 2.14 times higher for males than for females. Accordingly, q_1 varies with the proportion of workers over 35 who are male. In addition, data from Blanchard and Diamond (1990) show that, relative to males between the ages of 25 and 34, separations are 4.27 times higher for males between ages 16 and 24, 4.58 times higher for females between ages 16 and 24, and 2.48

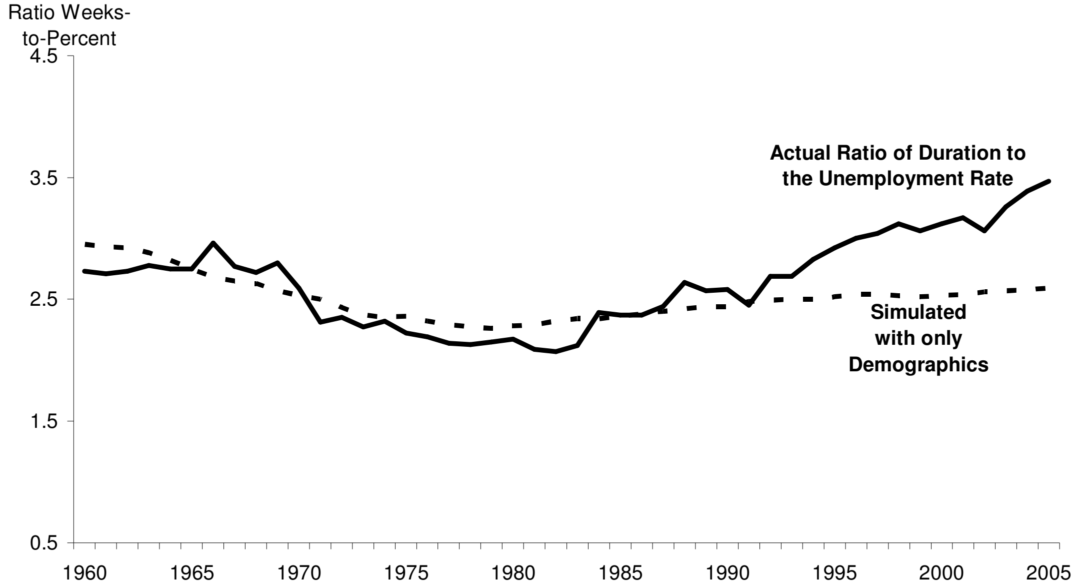
times higher for females between ages 25 and 34. Thus, the value of q_2 was adjusted to account for the proportion of young workers who fall into each of these subcategories.

For each group of workers (i.e., older and younger), the unemployment rate in each year was calculated by subtracting the value of q in that year from the baseline value of q. Then, this difference is multiplied by the value of du/dq from equation 7, and the product is added to the baseline value of the unemployment rate. However, as previously discussed, changes in q within an age group do not affect the average duration of unemployment for that age group.

Table 1 reports the results of a simulation in which the values of q_1 and q_2 are determined solely by the demographic variables (i.e., the age and gender composition of the workforce). This table lists the simulated values of duration, the natural rate, and the ratio between duration and unemployment, along with the actual ratio and the CBO's estimates of the NAIRU. Figure 3 plots the actual ratio of duration to the unemployment rate with the demographically simulated ratio over the sample period, and Figure 4 shows the simulated and CBO-estimated natural rates. Between 1960 and 1991, the demographic simulation does a very good job of predicting the natural rate and does a reasonably good job of predicting the duration-unemployment ratio. However, while the simulated data match actual data reasonably well between 1960 and 1991, after 1991 the demographic-simulation model substantially under-predicts the duration-unemployment ratio and substantially over-predicts the CBO's natural rate.

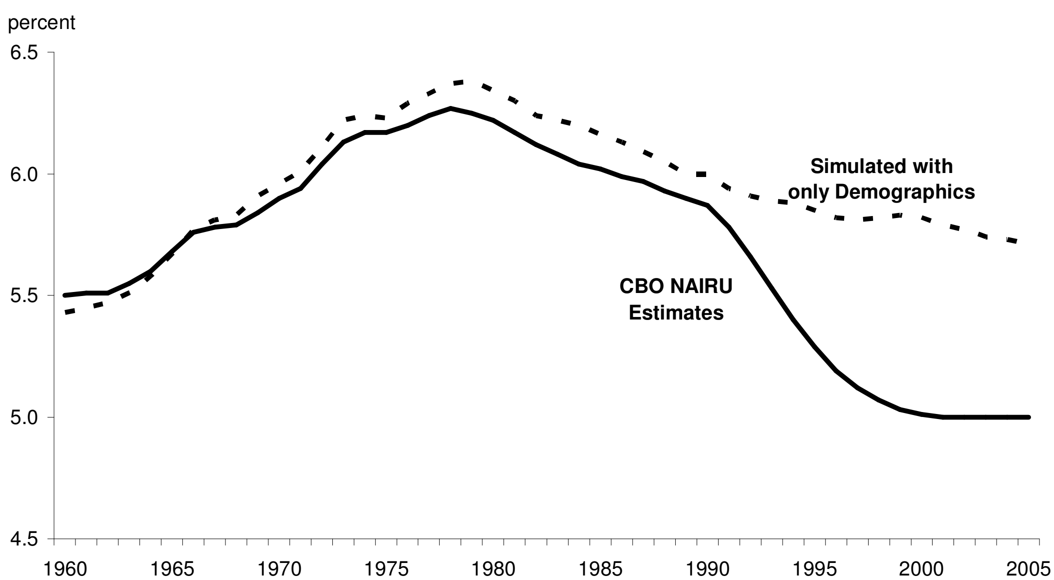
A plausible explanation for the rise in the duration-unemployment ratio and the fall in the natural rate is that there was a decline in the separation rate after 1991, as reflected in Shimer's (2005) estimates of the rate of job separation.[5] Between 1960 and 1990, the demographically-based simulation of q does a reasonably good job of

Figure 3: Demographics Alone Cannot Account for the Relative Rise of Duration to the Unemployment Rate



Sources: Bureau of Labor Statistics and authors' calculations

Figure 4: Demographics Alone Cannot Track the Decline in CBO's NAIRU Estimates during the 1990s

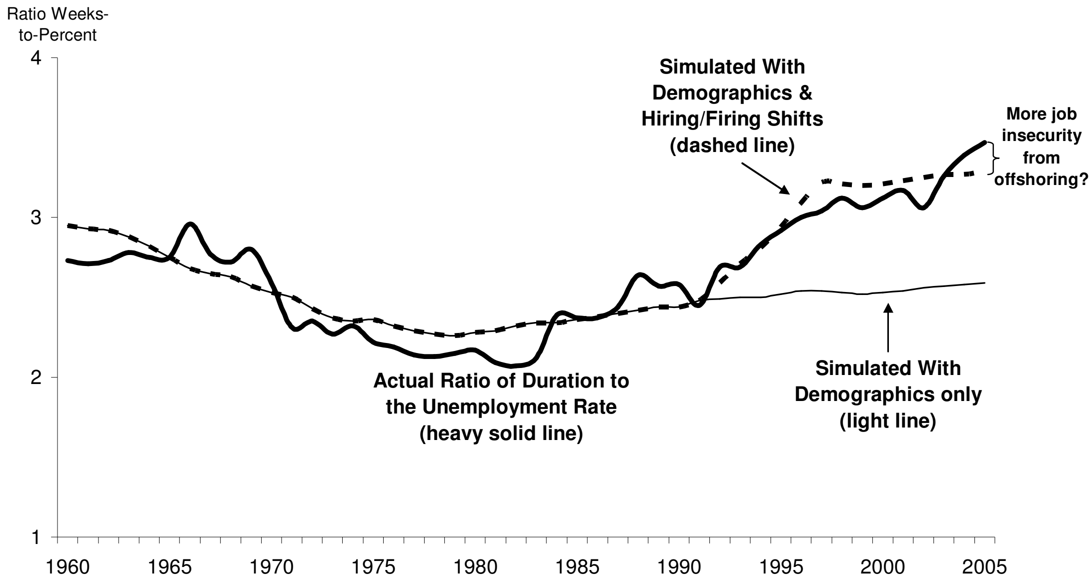


Sources: Congressional Budget Office (CBO) and authors' matching Shimer's estimates of the actual separation rate. For example, from 1960-70, the simulated value of q rose 13% and Shimer's estimates, by 11%. Between 1970 and 1980, the simulated value rose 10% and Shimer's estimate increased 12%. Then, from 1980-90, the simulated value declined by 8% and Shimer's estimate fell by 13%. However, between 1991 and 1997, Shimer's estimate of the separation rate fell much more (-16.5%) than the demographically-based simulated value (-4%). These patterns suggest that factors other than demographics lowered the separation rate in the 1990s.

To examine the effects of these changes on the duration-unemployment ratio and the natural rate of unemployment, it is assumed that the separation rate falls 2% a year between 1992 and 1997, in addition to changes stemming from demographics. Thus, the separation rate in 1997 is assumed to be 12% lower than it otherwise would have been. Table 2 shows how simulation results are altered when q is adjusted in this way. Based on simulations incorporating demographic changes and the above shift in the separation function, Figures 5 and 6 plot, respectively, the simulated and actual duration-unemployment ratios and the simulated and the CBO natural rates. As illustrated by these figures, simulations of the ratio and NAIRU much more closely track the actual ratio and natural rate from 1992-2005 when simulations reflect both demographic factors and plausible shifts in the separation function, rather than demographic factors alone.

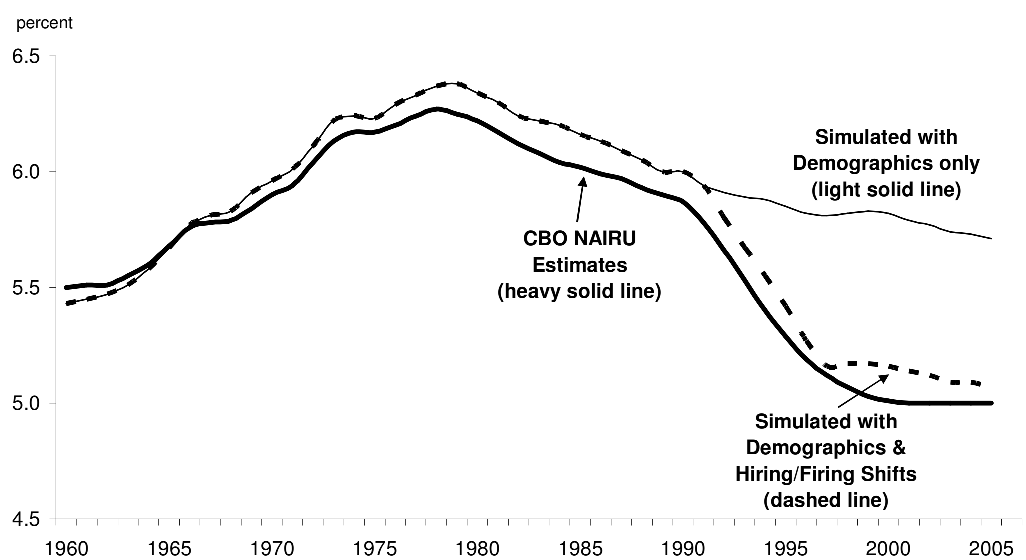
A more recent rise in the duration-unemployment ratio since 2004 is consistent with the view that increased outsourcing and globalization may have more recently increased job insecurity, suggesting that another shift in the separation function may be occurring. Nevertheless, this study does not try to pinpoint the precise source of shifts in separation practices, nor do the techniques employed lend themselves to doing so.

Figure 5: Demographics and Shifts in Turnover Behavior Track the mid-1990s' Relative Rise of Duration to the Unemployment Rate



Sources: Bureau of Labor Statistics and authors' calculations.

Figure 6: Demographics and Shifts in Turnover Behavior Help Track CBO NAIRU Estimates in the 1990s



Sources: CBO and authors' calculations.

III. The Estimated Impact of Higher, Relative Duration on Inflation

The simulation results presented earlier illustrate how the combination of the aging of the labor force and a shift in hiring and firing practices could provide an internally consistent explanation for the behavior of the natural rate and average duration of unemployment. To complement these findings, we assess whether our hypothesis is externally consistent using more traditional estimation techniques. In particular, we test whether the ratio of duration to the unemployment rate adds marginal information to expectations-augmented Phillips Curve or NAIRU models, and how the inclusion of the duration ratio affects the NAIRU estimates and overall performance of this framework.

Specification and Variables

According to the NAIRU framework popularized by Gordon (1977) and based on insights from Friedman (1968) and Phelps (1967, 1968), inflation can be modeled as:

$$\pi_t = \alpha_0 E_{t-1}(\pi_t) + \gamma(U_t^* - U_t), \tag{5}$$

where E is the expectations operator, α_0 is constrained to equal 1, U is the civilian unemployment rate, and U^* is the NAIRU. In practice, an energy price shock term is added to control for the effect of supply shocks on the NAIRU and empirical proxies (usually lagged inflation or survey data on expectations) are used to control for inflation expectations. Although U^* is not directly observed, if a constant is added the NAIRU can be estimated from the following baseline model, which largely follows Fuhrer (1995):

$$\begin{aligned}\pi_t = \beta_0 + \beta_1 U_{t-1} + \sum_{i=1}^{12} \alpha_i \pi_{t-i} + \beta_2 ENERGY_{t-1} + \delta_0 NIXON_t \\ + \delta_1 NIXOFF_t + \beta_3 \Delta RER8Q_{t-2}\end{aligned} \tag{6}$$

where π= inflation measured by the core PCE deflator, most variables are lagged to avoid simultaneity bias, *ENERGY* is the 8-quarter growth rate of the ratio of PCE energy prices to the core PCE index, and *NIXON* and *NIXOFF* are the dummy variables to control for the effects of imposing and lifting the wage-price controls during the Nixon administration.[6] For internal consistency, the NAIRU specification constrains the sum of coefficients on lagged inflation to equal 1 since these lags jointly proxy for expected inflation. In equilibrium, inflation equals its expectation, implying that $U^* = -\beta_0 / \beta_1$.

Owing to biases in measuring inflation with the CPI (see Boskin, et. al., 1996) and measurement changes to the CPI that make it inconsistent over time (see p. 94, Council of Economic Advisors, 1999), core inflation is measured with the PCE deflator.[7] One minor difference from Fuhrer's specification is that eq. (6) omits the lagged change in the unemployment rate which tracks speed effects (changes in unemployment) because this variable is very insignificant in core PCE and wage inflation models, in contrast to core CPI models. A second difference is that the t-1 lag of the 8-quarter percent change in relative PCE energy prices (ENERGY) is used instead of the time t percent change in the PPI price of oil relative to the PPI used by Fuhrer, because the latter is highly insignificant. The longer period over which relative energy price changes is measured allows for longer pass-through effects and the t-1 lag avoids simultaneity concerns. Another minor difference is the inclusion of the t-2 lag of the 8-quarter growth rate of the real value of the dollar ($\triangle RER8Q_{t-2}$) as measured by the Federal Reserve Board's broadly defined weighted average series. The t-2 lag of this term fit better than the t or t-1 lags, likely reflecting delays in the pass through of exchange rate changes to retail prices.

The wage inflation specification models nominal wage inflation in the nonfarm business sector (π^w), and is similar to eq. (6) except that it includes lags of inflation measured with the implicit price deflator for non-farm business prices (π_t^{nf}) rather than core inflation and, to control for normal real wage increases, includes non-farm productivity growth over the prior 12 quarters (*PROD*12)[8]:

$$\pi_t^w = \beta_0 + \beta_1 U_{t-1} + \beta_2 \Delta U_{t-1} + \sum_{i=1}^{12} \alpha_i \pi_{t-i}^{nf} + \beta_3 ENERGY_{t-1} + \beta_5 PROD12_{t-1} + \delta_0 NIXON_t + \delta_1 NIXOFF_t + \delta_2 PROFSHAR_t \quad (7)$$

Eq. (7) also includes a variable (PROFSHAR) to control for large swings in compensation surrounding the exercise of previously earned stock options which can cause large and hard to predict swings in compensation growth. Because stock options are only tracked by the compensation series when exercised rather than when earned, controlling for these large swings is needed to avoid serial correlation and misspecification problems. For example, measured compensation growth surged in 2000:q1 because employees exercised many stock options near the stock market peak and compensation growth became negative in the following quarter. To control for large swings, PROFSHAR is defined to equal the gap between compensation growth and ECI private worker compensation growth, when the gap is at least 0.5 percent at a quarterly rate, and 0 otherwise. (The ECI series does not yet include the value of stock options either when earned or exercised.) Prior to the late 1990s, there are very few instances, reflecting that stock options are a relatively new phenomenon. Values of PROFSHAR are set equal to 0 before 1980:2 because the ECI data start in 1980:1 and this period was likely unaffected by stock option payments to any noticeable extent.

To assess whether duration adds marginal information, the ratio of duration to the unemployment rate is added to the baseline models in (6) and (7):

$$\begin{aligned}\pi_t = \beta_0 + \beta_1 U_{t-1} + \sum_{i=1}^{12}\alpha_i \pi_{g-i} + \beta_2 ENERGY_{t-1} + \delta_0 NIXON_t \\ + \delta_1 NIXOFF_t + \beta_3 \Delta RER8Q_{t-2} + \beta_4 X_{t-1}\end{aligned} \tag{8}$$

$$\begin{aligned}\pi_t^w = \beta_0 + \beta_1 U_{t-1} + \beta_2 \Delta U_{t-1} + \sum_{i=1}^{12}\alpha_i \pi_{t-i}^{nf} + \beta_2 ENERGY_{t-1} + \beta_5 PROD12_{t-1} \\ + \beta_3 \Delta RER8Q_{t-2} + \beta_4 X_{t-1} + \delta_0 NIXON_t + \delta_1 NIXOFF_t + \delta_2 PROFSHAR_t\end{aligned} \tag{9}$$

where X is a vector that can contain duration and/or demographic variables and β_4 can be a row vector of more than one column when duration and demographic variables are included. In this case, the NAIRU is not a constant and $U^* = -(\beta_0 + \beta_4 X_{t-1})/\beta_1$.

Two considerations about the form of the variable (*DURRAT*) are noteworthy. First, a ratio is used to help identify the extra information in duration because the unemployment rate and the average duration of unemployment are collinear. Second, duration tends to lag the unemployment rate by two quarters, which makes intuitive sense since unemployment usually rises first in recessions and the average length of unemployment spells typically lengthens during the course of a recession until job creation resumes.[9] Two versions of the ratio of duration to unemployment are used. The first version is the one-quarter lag of ratio of duration in weeks at time t to the unemployment rate lagged by two quarters (*DURRAT*), reflecting that duration lags the unemployment rate by two quarters. The second is the smoothed version of the first using a Hodrick-Prescott filter (*DURRATHP*, q=1600). The first version is the one-quarter lag of ratio of duration in weeks at time t to the unemployment rate lagged by two quarters (*DURRAT*), reflecting that duration lags the unemployment rate by two quarters.

The main advantage of *DURRAT* relative to *DURRATHP* is that the marginal information in duration may be better identified because short-run movements in duration are better captured in *DURRAT*. The disadvantage of *DURRAT* is that it is noisier than *DURRATHP* since *DURRAT* displays some short-term swings that follow short-run changes in unemployment. Consequently, *DURRAT* yields noisier NAIRU estimates than does *DURRATHP*.

The unemployment rate and the average length of duration used as variables or in the construction of variables are adjusted for the 1994 changes in the household employment survey. Pre-1994 levels of the unemployment rate are adjusted upward by a multiplicative factor of 1.009 and average duration is adjusted upward by a multiplicative factor according to estimates based on overlapping data by Polivka and Miller (1998).[10]

Two approaches are used to control for demographic shifts. First, in some regressions which use the overall unemployment rate, the proportion of the labor force 35 years or older is added as an explanatory variable. Second, we replace the non-interacted unemployment rate with a demographically adjusted rate using a procedure similar to that of Shimer (1998, 2001). In particular, our demographically adjusted unemployment rate equals the weighted average sum of the unemployment rate of different age groups at time t multiplied by each group's share of the labor force during all of 1980.[11]

Regression Results

Regression results for core PCE and nominal wage inflation are presented in Tables 3 and 4, respectively. In each table, two sets of regressions are presented. In the first set, there are six models using the unemployment rate, with the baseline model shown in column 1 and with models 2 and 3 adding the duration-unemployment ratio (*DURRAT*) and the demographic ratio (*AGE*35+) to the baseline model, respectively. Model 4 adds both variables to the baseline model, while models 5 and 6 replace *DURRAT* in models 2 and 4 with *DURRATHP*. The second set of models replaces the non-interactive unemployment rate with the demographically adjusted rate, with models 7 and 8 corresponding to models 1 and 2. Owing to the use of 12 lags of inflation and the availability of core PCE and non-farm wage data since 1959, regressions are estimated over a common sample of 1962:q2-2005:q4. Consistent with the NAIRU approach, the constants and coefficients on the level of unemployment are statistically significant. The energy variable is statistically significant in most core PCE inflation models, but is insignificant in each wage model, consistent with the plausible case that opposing effects of energy shocks on labor supply and demand may result in an ambiguous net effect of energy shocks on wage inflation. Medium-run productivity growth and the variable controlling for stock options are statistically significant in each nominal wage regression.

Several notable patterns emerge across the tables. First, the duration-unemployment ratio, DURRAT, is always at least marginally statistically significant in the core PCE and wage models, while the ratio smoothed by the Hodrick-Prescott filter is not significant in the presence of the even less significant AGE35+ variable. Second, with respect to the core PCE models, the separate demographic variable (AGE35+) is significant only in the absence of the duration ratio (models 3 versus 4 and 6 in Table 3), whereas the unsmoothed duration ratio is significant, albeit to a lesser degree, in the presence of AGE35+. In the wage models, the duration ratios are still, albeit to a lesser extent, statistically significant in the presence of AGE35+, which is not significant in the presence of either duration variable. Third, in models using the demographically-adjusted unemployment rate, the duration ratio is significant in regressions of core PCE and non-farm wage (model 8 in Tables 3 and 4). Fourth, across corresponding models, the duration ratio (*DURRAT*) has a smaller t-statistic in the presence of AGE35+ (models 2 versus 4 in each table) or in models using the demographically adjusted unemployment rate (models 2 versus 8 in each table). This pattern plausibly reflects that movements in the duration ratio reflect demographic trends and other factors (e.g., shifts in hiring or firing behavior), consistent with the simulation results. Fifth, in the price inflation models 2-6 the coefficients and variable values imply that the NAIRU was between 4.3 and 5.0 percent in 2005:Q4, well below the baseline model 1 estimate of 5.84 percent. NAIRU estimates from the *DURRAT* and *DURRATHP* models imply that the NAIRU fell sharply in the 1990s, in contrast to the fixed, baseline model estimate (Figure 7).

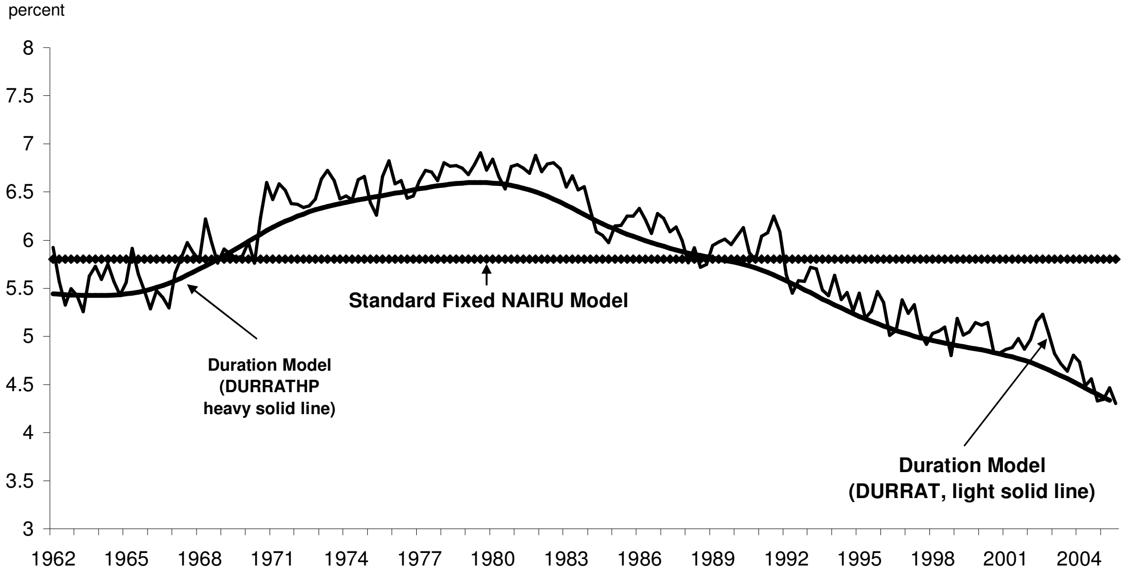
Sixth, this pattern arises using the demographically adjusted unemployment rate, with the NAIRU estimate from baseline price model (model 7, Table 3) at 6.2 percent and that from the duration model (model 8) at a lower 5.3 in 2005:Q4. [In 2005:Q4, the demographically adjusted rate, 5.8, exceeded the official rate of 5.0%.] Seventh, there is evidence of short-term serial correlation in residuals in wage models that omit either the duration or AGE35+ variables, suggesting that the standard models (model 1) or the simple demographically adjusted unemployment model (model 7) are miss-specified. Finally, using the overall unemployment rate, the NAIRUs at the end of 2005 from the wage equations (models 2-6) were notably lower when the duration ratio is included, in a range between 4.8 to 5.2 percent, versus 7.6 percent in the baseline model (model 1).

The improved performance of NAIRU models when adding this ratio parallels the simulation model results in an important aspect. In particular, demographic shifts could only partially account for the rise of the observed duration ratio in the 1990s in the simulation models, which implied a possible role for other factors, such as lower job security arising from shifts in the labor hiring or firing functions. This finding is consistent with the greater significance of the duration ratio than the demographic variable in the NAIRU models of price inflation. Also noteworthy, is that calibration experiments yield NAIRU and duration ratio estimates that are reasonably similar to the observed duration ratio (recall Figure 5) and the NAIRU results implied by the duration-modified model (model 2) of core PCE inflation (compare Figures 6 and 7).

IV. Conclusion

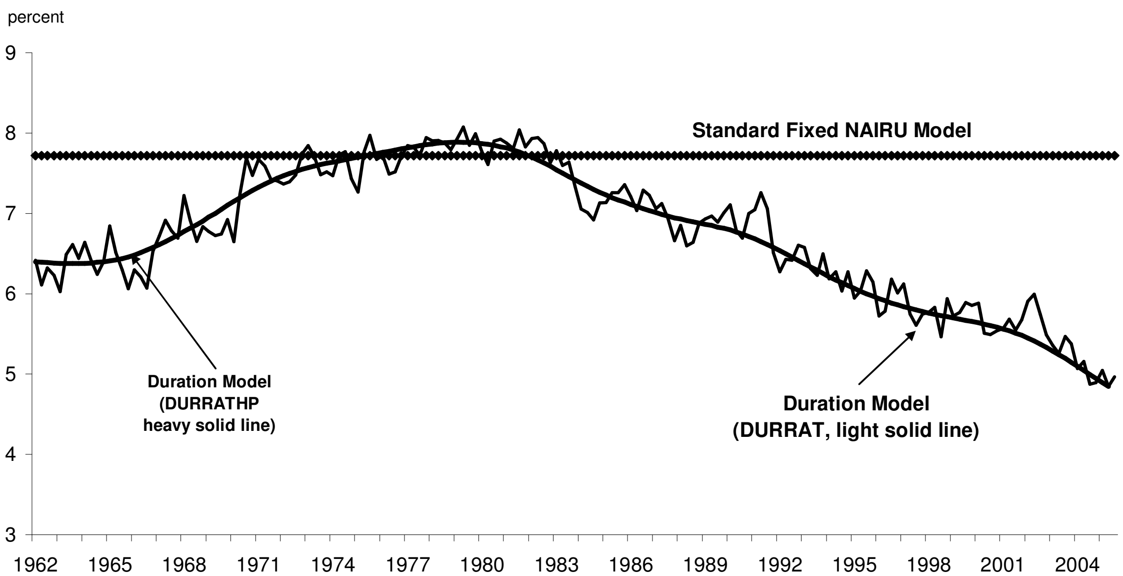
Since the mid 1980s, two major macro-labor indicators have shifted substantially, with the natural unemployment rate falling and the ratio of duration to the unemployment rate rising. Our simulation and regression results attribute these developments to a combination of the aging of the baby boom generation and a decline in job turnover, the latter of which may plausibly stem from decreased job security. By using both approaches, we test the internal and external consistency of the view that both factors have played important roles. In addition, other evidence suggests that declines in job turnover stemming from reduced voluntary quits may have been induced by increased job insecurity, consistent with survey data on worker perceptions of job insecurity and a greater relative tendency for unemployment to arise from dismissals than quits.

Figure 7: In Core PCE Models, the NAIRU Varies with Duration



DURRATHP controls for the lead of unemployment over duration by equaling the ratio of average duration in quarter t o the time t-2 unemployment rate., using data adjusted for the 1994 survey break and run through a Hodrick-Prescott filter. DURRAT controls for the lead of unemployment over duration by equaling the ratio of average duration in quarter t to the unemployment rate at time t-2. Sources: Bureau of Labor Statistics and authors' calculations.

Figure 8: In Compensation Models, the NAIRU Varies with Duration



DURRATHP controls for the lead of unemployment over duration by equaling the ratio of average duration in quarter t o the time t-2 unemployment rate., using data adjusted for the 1994 survey break and run through a Hodrick-Prescott filter. DURRAT controls for the lead of unemployment over duration by equaling the ratio of average duration in quarter t to the unemployment rate at time t-2. Sources: Bureau of Labor Statistics and authors' calculations.

More specifically, in a calibration model based on an efficiency wage framework, the combination of an aging labor force and a fall in the number of workers leaving their jobs (due to greater insecurity) could account for the combination of the higher duration ratio and the lower NAIRU estimates observed since the late-1980s. As in Shimer (1998, 2001), we find that shifting demographics play an important role, but we argue, in contrast, that demographics are not sufficient to fully account for changes in the NAIRU. In this respect, our simulation and estimation results are consistent with those of Valletta (1999), who found that the firing-to-quit ratio has risen at a given level of the unemployment rate since the late-1980s, and accord with widespread anecdotal reports, Challenger data on layoffs, and survey evidence indicating that job security has declined. Our results are also consistent with Staiger, Stock, and Watson (2001), who found that demographic factors could not account for the apparent post-1992 decline in the NAIRU.

In line with our calibration results, the duration-unemployment ratio is highly significant in expectations-augmented Phillips curve models of price and wage inflation. Furthermore, regression results imply a sizable decline in the NAIRU during the 1990s, when traditional, time *in*variant NAIRU models broke down. Although our findings mainly provide an explanation for the poor performance of traditional NAIRU models in the 1990s, they also imply that marginal information might be gleaned from monitoring the duration of unemployment relative to the unemployment rate. Nevertheless, caution should be exercised in using the duration ratio as an additional indicator, since simulation results indicate that the duration ratio may not always move in lock step with the NAIRU.

In addition to the NAIRU results, although the ratio of duration to unemployment was also high in the 1950s and early 1960s before the baby boomers entered the labor force, much of the rise since the late 1980s appears to be linked to factors other than demographics. As shown earlier in Figure 3, additional evidence favoring a role for non-demographic factors is that the ratio of the mean duration of unemployment to the unemployment rate has risen by more than what is implied by historical relationships with the demographic composition of the labor force.

By drawing on both calibration and estimation techniques, our findings provide an internally and externally consistent explanation for the behavior of the duration ratio and an apparent decline in the NAIRU. In particular, results imply that the unemployment rate—with or without demographic adjustments—is not as useful an indicator of labor market slack because changes in labor practices can alter job turnover and job security in ways not fully or consistently reflected in the unemployment rate.[12] In this respect, our regression and simulation results are consistent with Milton Friedman's (1968, p.8) characterization of the "natural rate of unemployment" as, "not immutable and unchangeable. On the contrary, many of the market characteristics that determine its level are man-made and policy-made."

Nevertheless, duration can reflect extra information about job security in plausible general equilibrium simulation models, and in NAIRU models of inflation, duration has provided statistically and economically important information beyond that contained in the overall or demographically- adjusted unemployment rates. With Friedman's caveat in mind, additional information regarding the degree in labor market slack may be gleaned from monitoring relative movements in duration. Other changes in labor practices since the 1990s could further alter these relationships. For example, future extensions of our study could examine how intra- and inter-national outsourcing of services will affect measures of unemployment slack and their relationship to inflation.

In interpreting the labor markets of the last two decades, a combination of high unemployment and low job security temporarily gave rise to the "traumatized worker[13]" and unexpectedly large disinflation of the early 1990s, after which a still elevated perception of job insecurity accompanied low unemployment rates during the long boom of the 1990s and the weak economy of the early 2000s (consistent with Figure 2 and Valletta, 1999). A long-lasting shift in labor market practices has apparently allowed the economy to operate at lower overall unemployment rates nearer to 5 percent rather than to 6 percent. Indeed, there was only a mild acceleration in inflation when unemployment fell below 4.5 percent in the late 1990s, followed by a notable deceleration during the slow economy of the early 2000s when unemployment remained below 6.5%. For these reasons, after being temporarily "traumatized" in the early 1990s, workers appear to have remained "chastened" as evolving labor practices continued to threaten job security and deter them from seeking other jobs.

References

Abowd, John M. and Zellner, Arnold (1985), "Estimating Gross Labor Force Flows," *Journal of Business and Economic Statistics* 3, 254-83.

Abraham, Katharine G. and Shimer, Robert (2001), "Changes in Unemployment Duration and Labor Force Attachment," unpublished manuscript, August 2001.

Blanchard, Olivier J. and Diamond, Peter (1990), "The Cyclical Behavior of the Gross Flows of U.S. Workers," *Brookings Papers on Economic Activity*, 1990:2, 85-143.

Boskin, Michael J., et al. (1996), "Toward a more Accurate Measure of the Cost of Living," Final Report to Senate Finance Committee from Advisory Commission to Study the Consumer Price Index (December 4).

Council of Economic Advisors (1997), *Economic Report of the President*. Washington: U.S. Government Printing Office.

Council of Economic Advisors (1999), *Economic Report of the President*. Washington: U.S. Government Printing Office.

Duca, John V. (1996), "Inflation, Unemployment, and Duration," *Economics Letters* 52 (September), 293-98.

Duca, John V. and VanHoose, David D. (2001), "The Rise of Goods Market Competition and the Fall of Nominal Wage Contracting," *Journal of Macroeconomics* 23 (Winter), 1-29.

Friedman, Milton (1968), "The Role of Monetary Policy," *American Economic Review* 58 (March), 1-17.

Fuhrer, Jeffrey C. (1995), "The Phillips Curve is Alive and Well," *New England Economic Review of the Federal Reserve Bank of Boston* (March/April), 41-56.

Gali, Jordi (1992), "How Well Does the IS-LM Model Fit Postwar U.S. Data?" *Quarterly Journal of Economics* 107 (May), 709-38.

Gordon, Robert J. (1977), "Can the inflation of the 1970s be explained?" *Brookings Papers on Economic Activity* 1, 253-76.

Gordon, Robert J. (1997), "The time-varying NAIRU and its implications for economic policy," *Journal of Economic Perspectives* 11, 11-32.

Greenspan, Alan (2004), "Testimony of Chairman Alan Greenspan: The Economic Outlook," to the Joint Economic Committee, U.S. Senate, April 21, http://www.federalreserve.gov/boarddocs/testimony/2004/20040421/default.htm.

Hall, Robert E. (1982), "The Importance of Lifetime Jobs in the U.S. Economy," *American Economic Review* 72 (September), 716-24.

Hyclak, Thomas and Ohn, Jonathan (2002), "Wage Inflation and the Post-1991 Duration Puzzle," *Economics Letters* 73 (October), 97- .

Phelps, Edmund S. (1967), "Phillips Curves, Expectations of Inflation, and Optimal Unemployment Over Time," *Economica* 2 (3), 22-44.

Phelps, Edmund S. (1968), "Money-Wage Dynamics and Labor Market Equilibrium," *Journal of Political Economy* 76 (July/August), 678-711.

Polivka, Anne and Stephen Miller (1998), "The CPS After the Redesign," in *Labor Statistics Measurement Issues*, edited by John Haltiwanger, Marilyn Manser, and Robert Topel, pp. 249-289. University of Chicago Press, Chicago.

Rich, Robert and Tracy, Joseph (2004), "Uncertainty and Labor Contract Durations," *Review of Economics and Statistics* 86(1) (February), 270-87.

Shimer, Robert (1998), "Why is the U.S. Unemployment Rate So Much Lower?" <u>NBER Macroeconomics Annual</u>.

________ (2001), "The Impact of Young Workers on the Aggregate Labor Market," *Quarterly Journal of Economics* 116 (August), 969-1008.

________ (2005), "Reassessing the Ins and Outs of Unemployment," Working Paper, University of Chicago, June 2005.

Staiger, Douglas, Stock, James H., and Watson, Mark W. (1997), "The NAIRU, Unemployment, and Monetary Policy." *Journal of Economic Perspectives* 11: 33-50.

Staiger, Douglas, Stock, James H., and Watson, Mark W. (2001), "Prices, Wages, and the U.S. NAIRU in the 1990s," in: *The Roaring Nineties: Can Full Employment Be Sustainable?* Edited by Alan B. Krueger and Robert Solow, 3-60.

Stephens, Melvin, Jr. (2004), "Job Loss Expectations, Realizations, and Household Consumption Behavior," *Review of Economics and Statistics* 86(1) (February), 253-69.

Valletta, Robert (1999), "Declining Job Security," *Journal of Labor Economics* 17 (October), S170-S197.

Woodward, Bob (2000), *Maestro: Greenspan's Fed and the American Boom* (New York: Simon & Schuster).

Table 1: Simulated and Actual Values of the Natural Rate and the Duration-Unemployment Ratio (Baseline Simulations Including Only Demographic Factors)

Year	Simulated Duration	Simulated Natural Rate	CBO Natural Rate	Simulated Ratio	Actual Ratio
1960	15.76	5.43	5.5	2.9	2.73
1961	15.76	5.45	5.51	2.89	2.71
1962	15.76	5.47	5.51	2.88	2.73
1963	15.73	5.52	5.55	2.85	2.78
1964	15.69	5.59	5.6	2.81	2.75
1965	15.63	5.67	5.68	2.76	2.75
1966	15.57	5.76	5.76	2.7	2.96
1967	15.52	5.81	5.78	2.67	2.77
1968	15.49	5.83	5.79	2.66	2.72
1969	15.43	5.91	5.84	2.61	2.8
1970	15.36	5.97	5.9	2.57	2.59
1971	15.3	6.03	5.94	2.53	2.31
1972	15.19	6.13	6.04	2.48	2.35
1973	15.09	6.23	6.13	2.42	2.27
1974	15.04	6.26	6.17	2.4	2.32
1975	15.01	6.27	6.17	2.39	2.22
1976	14.97	6.32	6.2	2.37	2.19
1977	14.93	6.36	6.24	2.35	2.14
1978	14.91	6.39	6.27	2.34	2.13
1979	14.91	6.39	6.25	2.33	2.15
1980	14.92	6.37	6.22	2.34	2.17
1981	14.94	6.33	6.17	2.36	2.09
1982	14.98	6.28	6.12	2.39	2.07
1983	15.01	6.25	6.08	2.4	2.12
1984	15.04	6.21	6.04	2.42	2.39
1985	15.08	6.17	6.02	2.44	2.37
1986	15.12	6.13	5.99	2.47	2.37
1987	15.17	6.09	5.97	2.49	2.44
1988	15.23	6.05	5.93	2.52	2.64
1989	15.18	6.07	5.9	2.5	2.57
1990	15.32	6	5.87	2.55	2.58
1991	15.39	5.94	5.78	2.59	2.45
1992	15.45	5.91	5.66	2.62	2.69
1993	15.51	5.88	5.53	2.64	2.69
1994	15.56	5.87	5.4	2.65	2.83
1995	15.6	5.85	5.29	2.67	2.92
1996	15.67	5.81	5.19	2.7	3
1997	15.71	5.8	5.12	2.71	3.04
1998	15.73	5.81	5.07	2.71	3.12
1999	15.76	5.82	5.03	2.71	3.06
2000	15.79	5.8	5.01	2.72	3.12
2001	15.83	5.78	5	2.74	3.17
2002	15.86	5.76	5	2.75	3.06
2003	15.91	5.74	5	2.77	3.26
2004	15.92	5.73	5	2.78	3.39
2005	15.94	5.71	5	2.79	3.47

Table 2: Simulated and Actual Values of the Natural Rate and Duration-Unemployment Ratios (Simulated Shifts in Demographics & Turnover)

Year	Simulated Duration	Simulated Natural Rate	CBO Natural Rate	Simulated Ratio	Actual Ratio
1960	15.76	5.43	5.5	2.9	2.73
1961	15.76	5.45	5.51	2.89	2.71
1962	15.76	5.47	5.51	2.88	2.73
1963	15.73	5.52	5.55	2.85	2.78
1964	15.69	5.59	5.6	2.81	2.75
1965	15.63	5.67	5.68	2.76	2.75
1966	15.57	5.76	5.76	2.7	2.96
1967	15.52	5.81	5.78	2.67	2.77
1968	15.49	5.83	5.79	2.66	2.72
1969	15.43	5.91	5.84	2.61	2.8
1970	15.36	5.97	5.9	2.57	2.59
1971	15.3	6.03	5.94	2.53	2.31
1972	15.19	6.13	6.04	2.48	2.35
1973	15.09	6.23	6.13	2.42	2.27
1974	15.04	6.26	6.17	2.4	2.32
1975	15.01	6.27	6.17	2.39	2.22
1976	14.97	6.32	6.2	2.37	2.19
1977	14.93	6.36	6.24	2.35	2.14
1978	14.91	6.39	6.27	2.34	2.13
1979	14.91	6.39	6.25	2.33	2.15
1980	14.92	6.37	6.22	2.34	2.17
1981	14.94	6.33	6.17	2.36	2.09
1982	14.98	6.28	6.12	2.39	2.07
1983	15.01	6.25	6.08	2.4	2.12
1984	15.04	6.21	6.04	2.42	2.39
1985	15.08	6.17	6.02	2.44	2.37
1986	15.12	6.13	5.99	2.47	2.37
1987	15.17	6.09	5.97	2.49	2.44
1988	15.23	6.05	5.93	2.52	2.64
1989	15.18	6.07	5.9	2.5	2.57
1990	15.32	6	5.87	2.55	2.58
1991	15.39	5.94	5.78	2.59	2.45
1992	15.45	5.8	5.66	2.67	2.69
1993	15.5	5.66	5.53	2.74	2.69
1994	15.55	5.55	5.4	2.81	2.83
1995	15.59	5.41	5.29	2.88	2.92
1996	15.66	5.27	5.19	2.97	3
1997	15.7	5.15	5.12	3.05	3.04
1998	15.72	5.16	5.07	3.05	3.12
1999	15.75	5.17	5.03	3.05	3.06
2000	15.78	5.15	5.01	3.06	3.12
2001	15.82	5.13	5	3.08	3.17
2002	15.85	5.12	5	3.1	3.06
2003	15.9	5.09	5	3.12	3.26
2004	15.91	5.08	5	3.13	3.39
2005	15.93	5.07	5	3.14	3.47

Table 3: Core PCE Inflation Regressions with Real Ex. Rate, Sample: 1962:Q2-2005:Q4[1]

	Overall Civilian Unemployment Rate						Demo. Adj. Unemp. Rate#	
Variable	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8
constant	0.0032	0.0082^{*}	0.0100^{*}	0.0082^{*}	0.0078^{**}	0.0083^{*}	0.0039^{**}	0.0069^{*}
	(4.81)	(4.57)	(3.30)	(2.59)	(4.27)	(2.55)	(5.38)	(4.23)
U_{t-1}	-0.0005^{**}	-0.0008^{**}	-0.0008^{**}	-0.0008^{**}	-0.0007^{**}	-0.0008^{**}	-0.0006^{**}	-0.0007^{**}
	(-5.02)	(-5.93)	(-5.22)	(-5.11)	(-5.74)	(-4.98)	(-5.57)	(-5.7)
$NIXON_t$	-0.0040^{**}	-0.0047^{**}	-0.0040^{**}	-0.0047^{**}	-0.0044^{**}	-0.0044^{**}	-0.0041^{**}	-0.0045^{**}
	(-2.95)	(-3.47)	(-2.96)	(-3.38)	(-3.27)	(-3.18)	(-3.08)	(-3.36)
$NIXOFF_t$	0.0110^{**}	0.0107^{**}	0.0108^{**}	0.0107^{**}	0.0107^{**}	0.0107^{**}	0.0109^{**}	0.0104^{**}
	(7.21)	(7.19)	(7.16)	(7.17)	(7.12)	(7.10)	(7.28)	(6.97)
$ENERGY_t$	-0.0003	-0.0002	-0.0003	-0.0002	-0.0001	-0.0001	-0.0003	-0.0004
	(-0.24)	(-0.18)	(-0.19)	(-0.18)	(-0.09)	(-0.10)	(-0.22)	(-0.34)
$\Delta RER8Q_{t-2}$	-0.0045^{*}	-0.0043^{*}	-0.0037^{+}	-0.0043^{*}	-0.0042^{*}	-0.0041^{*}	-0.0045^{*}	-0.0067^{*}
	(-2.26)	(-2.24)	(-1.84)	(-2.17)	(-2.16)	(-2.05)	(-2.32)	(-2.07)
$DURRAT_{t-1}$		-0.0014^{**}		-0.0014^{+}				-0.0008^{*}
		(-3.01)		(-1.89)				(-2.03)
$DURRATHP_{t-1}$					-0.0013^{**}	-0.0012		
					(-2.71)	(-1.40)		
$AGE35+_{t-1}$			-0.00010^{*}	0.000001		-0.0001		
			(-2.31)	(0.01)		(-0.18)		
NAIRU, 05:Q4	5.84%	4.39%	4.99%	5.08%	4.39%	4.44%	6.22%#	5.29%#
LM(1)	1.03	1.61	1.46	1.79	1.97	1.94	1.71	1.97
LM(2)	1.50	1.81	1.68	1.81	2.04	2.02	1.82	1.99
q(24)	19.14	15.57	15.76	15.58	16.33	16.14	17.88	16.88
R^2	.8854	.8910	.8885	.8902	.8898	.8892	.8889	.8911

1. Sums of coefficients for lags of inflation not reported as the sum is constrained to = 1. *(**,$^{+}$): significant at the 5% (1%, 10%) level.

Table 4: Nominal Wage Inflation Regressions With Real Ex. Rate (NonFarm Business Sector), Sample: 1962:Q2-2005:Q4[1]

	Overall Civilian Unemployment Rate						Demo.-Adj. Unemployment#	
Variable	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8
constant	0.0124**	0.0269**	0.0363**	0.0278**	0.0271**	0.0269**	0.0145**	0.0229**
	(6.35)	(8.27)	(5.94)	(4.29)	(8.10)	(4.02)	(7.49)	(7.93)
U_{t-1}	-0.0016**	-0.0022**	-0.0023**	-0.0022**	-0.00215**	-0.0021**	-0.0019**	-0.0020**
	(-6.39)	(-8.49)	(-7.84)	(-7.41)	(-8.42)	(-7.20)	(-7.58)	(-8.48)
$NIXON_t$	-0.0042	-0.0060*	-0.0037	-0.0059*	-0.0053*	-0.0053*	-0.0045+	-0.0057*
	(-1.57)	(-2.41)	(-1.44)	(-2.29)	(-2.13)	(-2.09)	(-1.75)	(-2.28)
$NIXOFF_t$	0.0067*	0.0060*	0.0065*	0.0060*	0.0059*	0.0059*	0.0067*	0.0058*
	(2.21)	(2.12)	(2.25)	(2.12)	(2.08)	(2.08)	(2.31)	(2.07)
$ENERGY_t$	0.0009	0.0013	0.0013	0.0013	0.0018	0.0018	0.0010	0.0013
	(0.34)	(0.37)	(0.50)	(0.53)	(0.72)	(0.72)	(0.40)	(0.54)
$PROD12_{t-1}$	0.0008*	0.0014**	0.0016**	0.0014**	0.0014**	0.0014**	0.0008**	0.0013**
	(2.49)	(4.24)	(4.42)	(3.74)	(4.25)	(3.64)	(2.72)	(4.08)
$DURRAT_{t-1}$		-0.0047**		-0.0046*				-0.0031**
		(-5.36)		(-3.26)				(-3.79)
$DURRATHP_{t-1}$					-0.0049**	-0.0049**		
					(-5.23)	(-3.05)		
$AGE35+_{t-1}$			-0.0004**	-0.00002		-0.00006		
			(-4.11)	(-0.16)		(-0.04)		
$\triangle RER8Q_{t-2}$	-0.0062	-0.0053	-0.0026	-0.0052	-0.0048	-0.0049	-0.0064	-0.0057
	(-1.55)	(-1.44)	(-0.67)	(-1.33)	(-1.30)	(-1.25)	(-1.65)	(-1.53)
$PROFSHAR_t$	0.8887**	0.9367**	0.8912**	0.9351**	0.9351**	0.9355**	0.8926**	0.9335**
	(12.36)	(14.01)	(13.01)	(13.79)	(13.93)	(13.70)	(12.97)	(13.95)
NAIRU, 05:Q4	7.65%	4.87%	5.17%	4.85%	4.76%	4.76%	7.77%#	5.88%#
LM(1)	7.66**	1.00	2.54	1.00	1.31	1.32	3.65+	0.97
LM(2)	10.84**	1.19	3.24	1.20	1.54	1.56	5.02+	1.20
q(24)	49.12**	30.19	27.41	30.09	30.33	30.37	33.85	31.12
R^2	.7360	.7758	.7604	.7744	.7756	.7743	.7565	.7757

1. Sums of coefficients for lags of inflation not reported as the sum is constrained to = 1. *(**,+): significant at the 5% (1%, 10%) level.

Endnotes

[1] While the probability of hire should be related to the total separation rate into nonemployment (including both unemployment and out of the labor force), the hiring probability can be expressed as a function of q if the proportion of separations that are into unemployment (relative to total separations into nonemployment) remains relatively constant over time. The probability of hire can be expressed as a function either of separations into unemployment or of total separations into nonemployment. The choice to express the hiring probability as a function of the probability of separation into unemployment is made for expositional convenience.

[2] If an individual moves directly from one employer to another employer, the number of vacancies in the aggregate economy is not affected, so separations that occur because of job changes do not affect the probability of hire.

[3] Monthly separation rates are calculated from the equation $1-(1-q)^{4.35}$, since there are, on average, 4.35 weeks per month.

[4] This value of q, the probability of a separation into unemployment, is somewhat higher than the value estimated by Blanchard and Diamond (1990), who used data from the Current Population Survey (CPS), as adjusted by Abowd and Zellner (1985). According to Blanchard and Diamond, transitions between employment and unemployment averaged 1.29% of employment on a monthly basis. However, the figure reported in Blanchard and Diamond omits those who are separated from their jobs in one month, experience a short spell of unemployment, and are rehired before the next month's survey.

[5] These data are available at http://home.uchicago.edu/~shimer/data/flows/sep.dat.

[6] NIXON equals 1 during the first two quarters of price controls (1971:3-71:4) and 0, otherwise, while NIXOFF equals 1 during the first two quarters when price controls were no longer in effect (1974:2-74:3). These variables differ slightly from those of Gordon (1977), which were less statistically significant and whose inclusion did not eliminate serially correlated errors in many similar (mainly baseline) regressions.

[7] Results were similar using the core CPI, but some of the post-1994 drop in the NAIRU derived from CPI regressions may be an artifact of changes in CPI measurement methodology designed to reduce bias.

[8] Overall prices outperformed PCE prices reflecting that firms pay the marginal product of labor (productivity plus wages deflated by output prices) in the long-run (pp. 147-49 and 151, *Economic Report of the Presiden*t, 1997). A productivity variable (PROD12) was added in order for the wage equations to be well-behaved, as in Staiger, Stock, and Watson (2001). The span of the productivity term mirrors the 12 quarterly inflation lags without unduly reducing the degrees of freedom by including 12 noisy lags of quarterly productivity growth.

[9] This is consistent with the classification of the unemployment rate as a coincident economic indicator and the duration of unemployment as a lagging indicator by the Conference Board.

[10] To construct a multiplicative adjustment similar to that for the unemployment rate, we multiplied pre-1994 data by the ratio of average duration for new survey technique to the old (17.19/14.96) using figures computed by Miller and Polivka (1998).

[11] We could not construct a demographically-adjusted duration ratio because age-specific duration data are unavailable. The duration ratio tests whether duration adds marginal information aside from demographics.

[12] Indeed, after noting that many unemployed workers exhausted their unemployment benefits in early 2004, Federal Reserve Chairman Alan Greenspan (2004) testified to Congress that: "Moreover, the average duration of unemployment increased from twelve weeks in September 2000 to twenty weeks in March of this [2004] year. These developments have led to a notable rise in insecurity among workers."

[13] A coinage attributed to Alan Greenspan [see Woodward (2000, pp.168-69)].