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. We argue that these developments can be explained by a combination of demographics and other factors that have reduced job turnover. The internal consistency of this hypothesis is assessed through simulations that show the effect of separations on the NAIRU and the duration-unemployment ratio. We then test the external consistency of this hypothesis by adding the duration-unemployment ratio to a NAIRU model. Including this variable adds significant explanatory power in estimating the NAIRU and yields realistic estimates of the current NAIRU.

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 (Katz and Krueger, 1999, and 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 that is age 35 and older (Figure 1).

This study argues that both of these phenomena may have resulted from a fall in the job separation rate (i.e., separations out of employment). A decrease in separations means that there are fewer job vacancies at a given unemployment rate, implying that the hiring rate is lower at each unemployment rate and that the ratio of average duration to unemployment is consequently higher. In addition, both an efficiency wage model and a matching model predict that a fall in the separation rate leads to a fall in the natural rate.

Data from Shimer (2007) suggests that job separations have decreased more than predicted by demographics since the early 1990s. While it is difficult to pinpoint the exact reason for a decline in job turnover, there are several plausible explanations for this decrease. First, downsizing during the 1990-91 recession and its aftermath may have caused workers to perceive a lower degree of job security. Workers who are more worried about layoffs are probably less likely to change jobs, since they may be uncertain about the quality of the match with their new employer and since the most recently hired workers are often the first to be laid off. Consistent with this explanation, Figure 2 shows that the share of workers worried about their job security (based on surveys conducted by International Survey Research) has generally been higher since the late 1980s. Second, Abraham and Shimer (2001) find that women's labor force attachment has increased, as evidenced by a significant decrease in the transition rate from employment to out of the labor force. Third, selection problems in obtaining health benefits at a new job may have become more significant as real health care costs have risen, discouraging workers from seeking new job matches. Fourth, as documented by Acemoglu (2002), within-group wage inequality has significantly increased in recent decades. If workers are risk averse, this increased inequality reduces the expected utility of unemployed individuals, decreasing the propensity to shirk and to quit into unemployment.

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 job turnover. First, we use an efficiency wage framework to show that a decrease in the equilibrium separation rate lowers the NAIRU and raises the duration-to-unemployment ratio. (An appendix available from the authors demonstrates that similar predictions arise in a matching model.) Given the effects of separations on both variables, this model predicts that the elasticity of the NAIRU with respect to the duration-unemployment ratio equals –1. We then simulate the NAIRU and the duration-unemployment ratio over 1960-2005, under the assumption that changes in the equilibrium separation rate are caused solely by demographics. These simulations generally track the natural rate and the duration-unemployment ratio through the early 1990s. Thereafter, however, the simulations substantially over-predict the natural rate and substantially under-predict the duration-unemployment ratio. Starting in the early 1990s, Shimer's (2007) measure of the actual separation rate falls significantly relative to the demographically estimated separation rate used in the simulations. When the simulations are rerun under the assumption that the separation rate falls in line with Shimer's measure, the simulations track the NAIRU and the duration-unemployment ratio very closely since the early 1990s. This pattern suggests that, together, demographic and non-demographic factors affecting the separation rate can account for the fall in the natural rate and the rise in the duration-unemployment ratio in recent decades.

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 this variable is that the model predicts that the NAIRU depends on the equilibrium separation rate, and the duration-unemployment ratio is related to the separation rate but is less affected by business cycle fluctuations. The coefficient on the duration-unemployment ratio is always significant, and the estimated elasticity of the NAIRU with respect to the duration-unemployment ratio is close to –1, the elasticity predicted by the theoretical model. Using this estimated model, we construct estimates of a time-varying NAIRU, and our results suggest that the duration-unemployment ratio can reasonably account for changes in the NAIRU over time.

II. Simulations of the Natural Rate and Average Duration

Model

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, and we denote the employment of each type as L_1 and L_2.

We assume firms pay efficiency wages and that the efficiency of type i workers is

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

where W_i is the wage the firm pays, $\overline{W}_i$ is the economywide average wage rate for type i workers, and h_i is the probability of hire each period for an unemployed worker of type i. While efficiency wage models generally treat efficiency as a function of the unemployment rate, it is more reasonable to model it as a function of the hiring rate. In the shirking model of Shapiro and Stiglitz (1984), the cost of losing one's job depends on the probability of finding a new job if dismissed, which depends on the hiring rate rather than on the unemployment rate, *per se*. In the turnover cost models of Stiglitz (1974), Schlicht (1978), and Salop (1979), the propensity to quit should depend on the probability of finding another job. While many quits are into another job rather than into unemployment, it is likely that the probability of finding another job when employed is closely related to the probability of finding a job when unemployed. Thus, the probability of hire for an unemployed worker should affect both quits into unemployment and quits into a different job.[1]

The probability of an unemployed worker being hired depends on the number of matches (as a percentage of the labor force) between the unemployed and employers with vacancies divided by the unemployment rate. Let $m(u_i, v_i)$ represent the matching rate for type i workers (where u is the unemployment rate and v is the vacancy rate). In equilibrium, outflows from employment must equal inflows, which implies that $(1-u_i)q_i = m(u_i, v_i)$, where q_i is the separation probability. The probability of hire for an unemployed worker equals the matching rate divided by the unemployment rate:

$$h_i = \frac{q_i(1-u_i)}{u_i} \qquad \text{with } h_q > 0 \text{ and } h_u < 0.\text{[2]} \tag{2}$$

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

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

Accordingly, the profits of a typical firm are

$$\pi = Pf\left(e^1\left(\frac{W_1}{\overline{\overline{W}}_1}, h(u_1, q_1)\right)L_1, e^2\left(\frac{W_2}{\overline{\overline{W}}_2}, h(u_2, q_2)\right)L_2 \right) - W_1L_1 - W_2L_2. \tag{3}$$

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

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

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

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

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

At the point where $W_i = \overline{W}_i$, the condition in equation (6) determines the economy's natural rate of unemployment, which depends on the separation probability. 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 and totally differentiating (6). Accordingly,

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

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

$$h_q = \frac{1-u}{u}, \quad \text{and} \quad h_u = -\frac{q}{u^2},$$

so the elasticity of the natural rate with respect to the separation rate is

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

The average duration of unemployment can be calculated as follows (assuming there is no duration dependence). In each period, the number of type *i* workers who have undergone a separation and remain unemployed is

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

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

$$\sum_{k=1}^{\infty} k L_i q_i (1-h_i)^k = \sum_{k=1}^{\infty} \sum_{m=k}^{\infty} L_i q_i (1-h_i)^m = L_i q_i \frac{1-h_i}{h_i^2}.$$[3]

Thus, the average duration of unemployment for type *i* workers can be expressed as

$$D_i = \frac{1}{h_i} = \frac{u_i}{q_i(1-u_i)}. \tag{8}$$

To calculate the impact of the separation rate on average duration, we differentiate (8) with respect to u and q, yielding

$$dD_i = \frac{q_i du_i - u_i(1-u_i) dq_i}{q_i^2 (1-u_i)^2}. \tag{9}$$

Since q indirectly affects D through its effect on u, it is necessary to substitute (7) into (9) to find the total effect of q on D, which is

$$\frac{dD_i}{dq_i} = \frac{q_i \dfrac{u_i(1-u_i)}{q_i} - u_i(1-u_i)}{q_i^2(1-u_i)^2} = 0. \tag{10}$$

The fact that $dD/dq=0$ means that the indirect effect of q on D exactly offsets the direct effect. 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.

The effect of the separation rate on the duration-unemployment ratio is

$$\frac{d(D_i/u_i)}{dq_i} = \frac{u_i \dfrac{dD_i}{dq_i} - D_i \dfrac{du_i}{dq_i}}{u_i^2}.$$

From equations (7) and (10), $du/dq=(1-u)(u/q)$ and $dD/dq=0$. Thus the elasticity of the duration-unemployment ratio with respect to the separation rate can be expressed as

$$\frac{d(D_i/u_i)}{dq_i}\frac{q_i}{(D_i/u_i)} = -(1-u_i). \tag{11}$$

Since the elasticity of the natural rate with respect to separations is $(1-u)$ and the elasticity of the duration-unemployment ratio with respect to separations is $-(1-u)$, the model predicts that the elasticity of the NAIRU with respect to the duration-unemployment ratio should equal –1.

The model developed in this subsection assumes that equilibrium unemployment is explained by an efficiency wage model. Another possibility is that unemployment arises from a matching model, as described in Pissarides (2000). While frictional unemployment is undoubtedly part of total unemployment, it probably does not explain all of unemployment. For example, Malcomson and Mavroeidis (2007) estimate a model in which long-term unemployment consists of a mixture of efficiency wage, frictional, and high wage unemployment. (High wage unemployment is defined as unemployment that occurs when wages exceed efficiency wages.) Their point estimate of the long-run unemployment rate in their sample is 5.9%, and they estimate that the shares of the various types of unemployment are 3.5% for efficiency wages, 1.7% for frictional, and 0.7% for high wages. These results suggest that efficiency wages are more important than labor market frictions in explaining the natural rate of unemployment.

The model developed in this subsection allows for frictional unemployment, since the probability of hire for an unemployed worker depends on a matching function that incorporates vacancies. In addition, it can be demonstrated that the matching model of Pissarides (2000) makes similar predictions about the effect of changes in the separation rate on the natural rate of unemployment. In an unpublished appendix the Pissarides model is calibrated, and it is demonstrated that the predicted elasticity of the natural rate with respect to the separation rate is almost identical in the matching model and the efficiency wage model.[4]

Calibration

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.[5] 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 relative baseline values of q, we make use of the fact that the average unemployment duration was 1.5 times longer for older workers than for younger workers. From equation (8), this condition implies that the separation rate is 3.44 times higher for younger workers than for older workers.

<u>Simulation results</u>

After determining baseline parameters 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 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 NAIRU in each year was calculated by finding the percentage difference between q and the baseline value of q

(based on the demographic composition of the group) and adjusting the NAIRU from its baseline value using the relationship in equation 7. However, as previously discussed, changes in q within an age group do not affect the average duration of unemployment for that age group.

Figures 3 and 4 show the results of simulations 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). 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 closely predicts the natural rate and reasonably tracks the duration-unemployment ratio. However, after 1991 the demographic simulation substantially under-predicts the duration-unemployment ratio and 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 a decline in the separation rate after 1991, as reflected in Shimer's (2007) estimates of the rate of job separation.[6] Figure 5 shows the separation rate predicted by demographics and Shimer's estimates of the actual separation rate, relative to their average values in 1960-64. Since the actual separation rate is countercyclical and fairly volatile, this figure also shows a cyclically adjusted and smoothed separation rate. To obtain this measure, the separation rate was regressed on a constant and the difference between the actual unemployment rate and the CBO's estimate of the natural rate, and a cyclically-adjusted separation rate was created by setting the unemployment rate equal to the natural rate in each year. The resulting series was then smoothed using a Hodrick-Prescott filter with a smoothing factor of 1600.

Between 1960 and the early 1990s the separation rate predicted by demographics closely tracks Shimer's measure of the actual separation rate. The unadjusted actual separation rate was below the demographically-predicted separation rate during the expansion in the late 1960s and was above it during the recession in the early 1980s, but these measures were close over most of this period. In addition, there is a very close relationship between the demographic separation rate and the cyclically adjusted and smoothed separation rate. Starting in the early 1990s, however, both the actual and the adjusted actual separation rates decline relative to the demographic separation rate. Between 1991 and 2005, the actual separation rate fell by 25.0%, while the separation rate predicted by demographics decreased by only 7.3%. Thus, the actual separation rate fell 19.1% more (1–(0.75/0.927)) than the decrease predicted by demographics, suggesting that factors other than demographics lowered the separation rate.

To examine how this "non-demographic" fall in the separation rate affects the natural rate of unemployment and the duration-unemployment ratio, the simulations were rerun under the assumption that the separation rate falls by a cumulative amount of 19.1% between 1991 and 2005 in even increments each year, in addition to changes stemming from demographics. These revised simulations are plotted in Figures 6 and 7, which show that simulations 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.

In efficiency wage models in which efficiency depends on the unemployment rate, the Solow (1979) condition determines the economy's natural rate. In contrast, the model developed in this section treats efficiency as a function of the hiring probability, which means that equation (6) determines the economy's equilibrium hiring probability.

Hiring data estimated by Robert Shimer suggests that, holding demographics constant, the hiring probability has been stable over time, barely changing from 0.428 in 1960 to 0.414 in 2005. In both these years, the economy was close to full employment, and the percentage of workers 35 and older was similar (62.8% in 1960 and 63.4% in 2005).[7]

While the hiring probability was stable over the 1960-2005 period, it is possible that future developments could cause a shift in the equilibrium hiring rate. If this occurs, then the model developed in this section would not exactly capture the relationship between the equilibrium separation rate and the natural rate.

The reason why the model in this section assumes there are two types of workers is that the hiring probability is much lower for older than for younger workers, as evidenced by the fact that average unemployment duration is 1.5 times longer for older workers. However, the assumption that there are two types of workers is not critical for obtaining the main results of this study. In a model having only one type of worker, simulations also show a large increase in the duration-unemployment ratio and a large decrease in the NAIRU in the post-1991 period.[8]

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 change in separation propensities 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

$$\pi_t = \alpha_0 E_{t-1}(\pi_t) + \gamma(U_t^* - U_t), \tag{12}$$

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):

$$\begin{aligned} \pi_t = \beta_0 + \beta_1 U_{t-1} + \sum_{i=1}^{12} \alpha_i \pi_{t-i} + \beta_2 ENERGY_{t-1} + \delta_0 NIXON_t \\ + \delta_1 NIXOFF_t + \beta_3 \Delta RER8Q_{t-2} \end{aligned} \tag{13}$$

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.[9] 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.[10] One minor difference from Fuhrer's specification is that eq. (13) 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 are 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 ($\Delta RER8Q_{t-2}$) 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. (13) 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 the quarterly average non-farm productivity growth over the prior 12 quarters (*PROD*12):[11]

$$\begin{aligned}\pi_t^w = \beta_0 + \beta_1 U_{t-1} + \beta_2 \Delta U_{t-1} + \sum_{i=1}^{12} \alpha_i \pi_{t-i}^{nf} + \beta_3 ENERGY_{t-1} + \beta_5 PROD12_{t-1} \\ + \delta_0 NIXON_t + \delta_1 NIXOFF_t + \delta_2 PROFSHAR_t.\end{aligned} \tag{14}$$

Eq. (14) 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:Q2 because the ECI data start in 1980:Q1 and this period was likely unaffected by stock option payments to any noticeable extent.

As discussed in Section II, the NAIRU should depend positively on the separation rate, so the NAIRU in period *t* can be expressed as $U_t^* = U_0^* + \mu(q_t - q_0)$. However, as demonstrated in Figure 5, the separation rate is significantly affected by business cycle fluctuations. A measure that is less cyclically sensitive is the duration-unemployment ratio, and Section II shows that the NAIRU is related to this ratio, since both the NAIRU and this ratio depend on the separation rate. Thus, the NAIRU can be expressed as $U_t^* = U_0^* - \psi[(D/u)_t - (D/u)_0]$, where ψ should equal 1 if the steady-state hiring probability remains constant over time.

To assess whether duration adds marginal information, the ratio of duration to the unemployment rate is added to the baseline models in (13) and (14):

$$\pi_t = \beta_0 + \beta_1 U_{t-1} + \sum_{i=1}^{12} \alpha_i \pi_{g-i} + \beta_2 ENERGY_{t-1} + \delta_0 NIXON_t \\ + \delta_1 NIXOFF_t + \beta_3 \Delta RER8Q_{t-2} + \beta_4 X_{t-1}, \tag{15}$$

$$\pi_t^w = \beta_0 + \beta_1 U_{t-1} + \beta_2 \Delta U_{t-1} + \sum_{i=1}^{12} \alpha_i \pi_{t-i}^{nf} + \beta_2 ENERGY_{t-1} + \beta_5 PROD12_{t-1} \\ + \beta_3 \Delta RER8Q_{t-2} + \beta_4 X_{t-1} + \delta_0 NIXON_t + \delta_1 NIXOFF_t + \delta_2 PROFSHAR_t \tag{16}$$

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 $U^* = -(\beta_0 + \beta_4 X_{t-1})/\beta_1$ using coefficients from the price equation. In the wage equation, the NAIRU also needs to reflect that normal real wage growth should rise one-for-one with productivity growth over the long-run, so that $\pi^w = PROD12$ + expected inflation when $U=U^*$. This implies that $U^* = -[\beta_0 + \beta_4 X_{t-1} - ((1-\beta_5)PROD12)]/\beta_1$ using coefficients from the wage equation. To make this adjustment easier to follow, *PROD*12 is defined as a decimal and is quarterly.[12] Also note that because the unemployment rate enters as a percent unlike the other variables and because the wage and inflation data are quarterly nonpercent rates, the magnitudes of the individual coefficients may differ from those in other studies. Nevertheless, by its construction, the implied NAIRU estimates and the fit of the equations are not affected by differences in scaling with other studies.

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.[13] Two versions of the ratio of duration to unemployment are used. The first version is the one-quarter lag of the 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 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.

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).[14]

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 rates of different age groups at time *t* multiplied by each group's share of the labor force during all of 1980.

Regression Results

Regression results for core PCE and nominal wage inflation are presented in Tables 1 and 2, 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:Q2-2006: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.

Also reported in these tables is the elasticity of the NAIRU with respect the duration-unemployment ratio ($\%\Delta U^{\bullet} / \%\Delta R$, where R is the duration-unemployment ratio). If the hiring rate remains stable over time, this elasticity should equal –1.

Several notable patterns emerge across the tables. First, the duration-unemployment ratio, *DURRAT*, is always statistically significant in the core PCE and wage models. This is also the case for the ratio smoothed by the Hodrick-Prescott filter except when it is marginally significant in the presence of the highly insignificant *AGE35+* variable. In addition, the elasticity of the NAIRU with respect to the duration-unemployment ratio lies between –0.79 and –1.00, and it is never statistically significantly different from –1, the value predicted by the model in Section II. Thus, the model developed in Section II accurately predicts the effect of the duration-unemployment ratio on the natural rate. 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 1), 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 wages (model 8 in Tables 1 and 2). 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 both 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 4.2 and 4.9 percent in 2006:Q4, well below the baseline model 1 fixed estimate of 5.75 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 8).

Sixth, this pattern arises using the demographically adjusted unemployment rate, with the NAIRU estimate from the baseline price model (model 7, Table 1) at 6.2 percent and that from the duration model (model 8) at a lower 5.3 in 2006:Q4. (In 2006:Q4, the demographically adjusted rate, 5.25, exceeded the official rate of 4.5%.) 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 model (model 1) and the simple demographically adjusted unemployment model (model 7) are misspecified. Finally, using the overall unemployment rate, the NAIRUs at the end of 2006 from the wage equations (models 2, 4, 5, and 6) were in a reasonable range between 4.0 and 4.3 percent when the duration ratio is included, unlike the bizarre -1.6 percent productivity adjusted rate from the baseline model (model 1) (which, before adjusting for productivity, yields a NAIRU of 7.8 percent.) A plausible interpretation of the latter unusual result is that it comes from a misspecified model which yields a smaller than true productivity coefficient, which works to lower the implied, productivity-adjusted NAIRU estimate.

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 affecting separations, such as lower job security. This finding is consistent with the greater significance of the duration ratio than the demographic variable in the NAIRU models of wage and price inflation. Also noteworthy is that calibration experiments yield duration ratio and NAIRU estimates that are reasonably similar to the observed duration ratio (recall Figure

6) and the NAIRU results implied by the duration-modified model (model 2) of core PCE inflation (compare Figures 7 and 8).

IV. Conclusion

Since the early 1990s, 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 an additional 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.

More specifically, in a calibration model based on an efficiency wage framework it is demonstrated that changes in the separation rate can explain much of the variation in the duration-unemployment ratio and the NAIRU since 1960. From 1960 until the early 1990s, the economy's separation rate was closely related to the demographic composition of the workforce. Since the early 1990s, however, the separation rate fell about 20% more than predicted by demographics, and a combination of an aging labor force and this additional decline in the separation rate can account for the combination of the higher duration ratio and the lower NAIRU estimates observed since the early 1990s. 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 the post-1991 decline in the NAIRU, consistent with the results of Katz and Krueger (1999) and Staiger, Stock, and Watson (2001).

While it is unclear what precisely caused the post-1991 decline in the separation rate, several possible explanations are discussed. One explanation that we find particularly promising is that the decreased separation rate may have resulted from factors leading to greater job insecurity faced by workers. Higher job insecurity is evidenced by surveys of workers indicating greater fear of layoffs, Challenger data on layoffs, and the results of Valletta (1999), who finds that the firing-to-quit ratio has risen at a given level of the unemployment rate since the late-1980s.

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.

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, our 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.[15] 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 provides 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 market behavior 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 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[16]" and unexpectedly large disinflation of the early 1990s. Afterwards, 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.

Appendix: Estimates of the Effect of Separation Probabilities on the Unemployment Rate in the Model of Pissarides (2000)
(Not intended for publication, available upon request from the authors)

Pissarides (2000) develops a model of the job matching, vacancies, and unemployment in a framework in which wages are determined by bargaining between firms and workers. As discussed on p. 18 of his book, the following three equations determine steady-state equilibrium in this model:

$$u = \frac{\lambda}{\lambda + \theta q(\theta)}, \tag{A1}$$

$$p - w - \frac{(r+\lambda)pc}{q(\theta)} = 0, \qquad \text{and} \tag{A2}$$

$$w = (1-\beta)z + \beta p(1 + c\theta), \tag{A3}$$

where λ represents the rate of flows into unemployment from employment (equivalent to the parameter q in Section II), θ is the ratio between vacancies and unemployment, $q(\theta)$ is the rate at which vacant jobs are filled, p represents the productivity of workers (i.e., the value of a job's output), w is the real wage, r is the interest rate, c is the real cost to firms of having a vacancy, β represents the relative bargaining power of workers, and z is a worker's real return when unemployed.[17] Pissarides defines $q(\theta)$ as $q(\theta) = m(u/v, 1)$. If it is assumed that the matching function can be expressed as $m = \gamma u^{a} v^{1-a}$, then $q(\theta) = \gamma\theta^{-a}$. With this specification for $q(\theta)$, (A1) can be rewritten as

$$u = \frac{\lambda}{\lambda + \gamma\theta^{1-a}}. \tag{A4}$$

Differentiating (A4) with respect to u, λ, and θ yields the following relationship between the unemployment rate, the separation rate, and the vacancy-unemployment rate:

$$du = \frac{\theta\gamma\theta^{-a}d\lambda - (1-a)\lambda\gamma\theta^{-a}d\theta}{[\lambda + \gamma\theta^{1-a}]^2}. \tag{A5}$$

To calculate the total effect of separations on unemployment, we need to calculate the effect of λ on θ. This effect can be calculated by the following procedure. We will assume that the worker's return when unemployed is a constant fraction, b, of the return when employed, so that $z = bw$. In addition, productivity will be normalized to equal 1. Under these assumptions, (A3) can be rewritten as

$$w = \frac{\beta}{1-(1-\beta)b}(1+c\theta). \tag{A6}$$

If (A6) is substituted into (A2), we obtain the relationship,

$$1 - \frac{\beta}{1-(1-\beta)b}(1+c\theta) - \frac{(r+\lambda)c}{\gamma\theta^{-a}} = 0. \tag{A7}$$

Differentiating (A7) with respect to λ and θ implies

$$-\frac{\beta}{1-(1-\beta)b}cd\theta - \frac{\gamma\theta^{-a}cd\lambda + a(r+\lambda)c\gamma\theta^{-a-1}d\theta}{\gamma\theta^{-a}} = 0, \text{ and}$$

$$\frac{d\theta}{d\lambda} = -\frac{\gamma c\theta^{-a}}{\frac{\beta c\gamma\theta^{-a}}{1-(1-\beta)b} + a(r+\lambda)c\gamma\theta^{-a-1}}. \tag{A8}$$

From (A5), the elasticity of the unemployment rate with respect to the separation rate is

$$\frac{du}{d\lambda}\frac{\lambda}{u} = \frac{\theta\gamma\theta^{-a} - (1-a)\lambda\gamma\theta^{-a}(d\theta/d\lambda)}{[\lambda + \gamma\theta^{1-a}]^2}\frac{\lambda}{u}, \tag{A9}$$

where $(d\theta/d\lambda)$ is given by (A8).

If values are assigned to the model's parameters, we can calculate the effect of the separation rate (λ) on the unemployment rate (u). The following assumptions are made about the model's parameters. First, β equals 0.5, meaning that firms and workers have equal bargaining power. Second, the value of b (the ratio between the returns to working and the returns to being unemployed) is assumed to equal 0.5. Third, in line with estimates of the average monthly separation rate in the United States in 1960-1970 (discussed in Section II), λ is set equal to 0.0146. Fourth, it is assumed that θ=0.3, based on the ratio of vacancies to unemployment in years in which the actual unemployment rate was within 0.5 of the Congressional Budget Office's (CBO) estimate of the natural rate.[18] Fifth, it is assumed that the real interest rate is 0.0033 per month, implying an annual real interest rate of 0.04. Sixth, in line with the average of the CBO's estimate of the natural rate for the 1960-70 period, the natural rate is assumed to equal 0.0567.

Values for a (the exponent in the matching function) have been estimated in previous studies, and Petrongolo and Pissarides (2001) report estimates of a obtained by several researchers. For the United States, these estimates generally are in the range of 0.35 to 0.50. Thus, calculations are performed both for a=0.35 and a=0.50. Given values for a, u, θ, and λ, the value of γ in the matching function is uniquely determined by (A4).

Substituting values of β, b, θ, λ, r, a, and γ into (A7) yields a solution for c. This value for c is then substituted into the expression for $(d\theta/d\lambda)$ in (A8), which is then substituted into (A9), yielding a value for the elasticity of the natural rate with respect to the separation rate. This elasticity is calculated to be 0.987 when a=0.35 and to equal 0.976 when a=0.50. In the efficiency wage model developed in Section II, a value of u=0.0567 means that the elasticity of the natural rate with respect to the separation rate is 0.943. Thus, a matching model and an efficiency wage model yield almost identical estimates of the elasticity of unemployment with respect to the separation rate.

Why the elasticities are almost identical in the two models can be explained by the effect of the separation ratio on the hiring probability. In the efficiency wage model, the equilibrium hiring probability is constant and thus is not affected by the separation probability. In the matching model, the hiring probability depends on the separation probability, but the effect is weak. Thus, both models predict similar elasticities of unemployment with respect to the separation rate.

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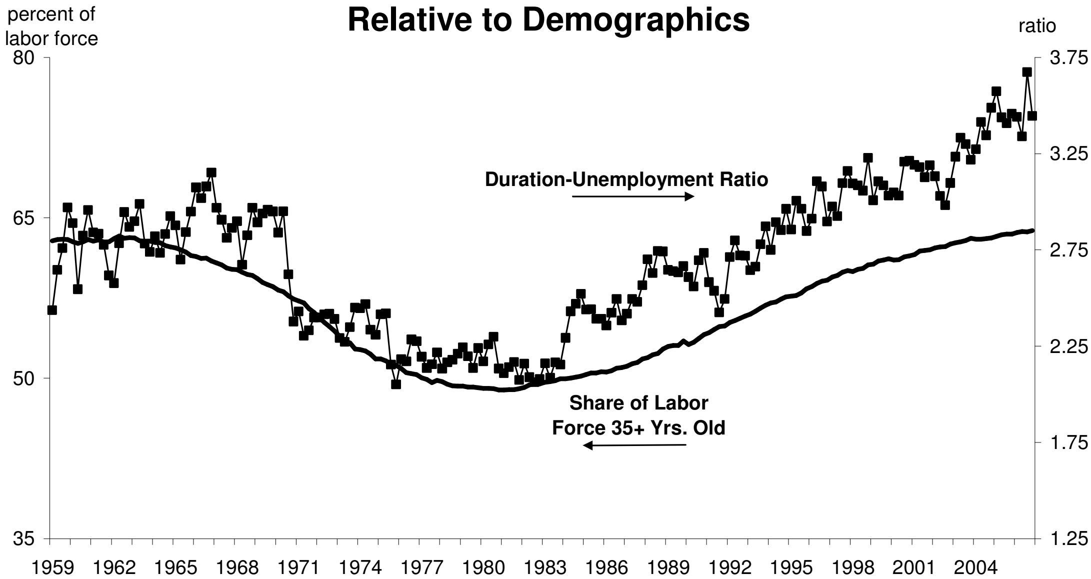
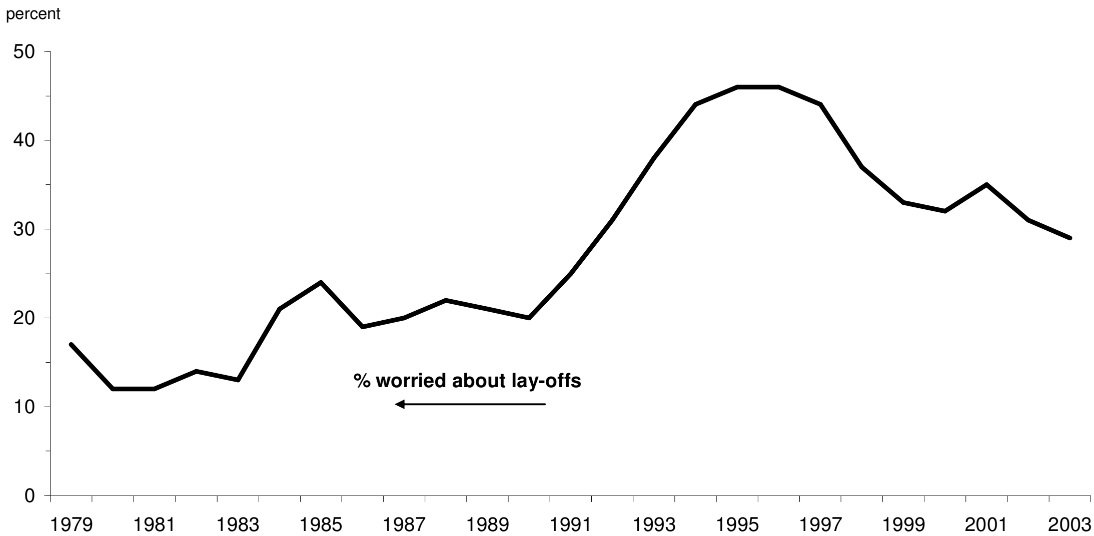
Figure 1: Duration-Unemployment Ratio Rises Relative to Demographics



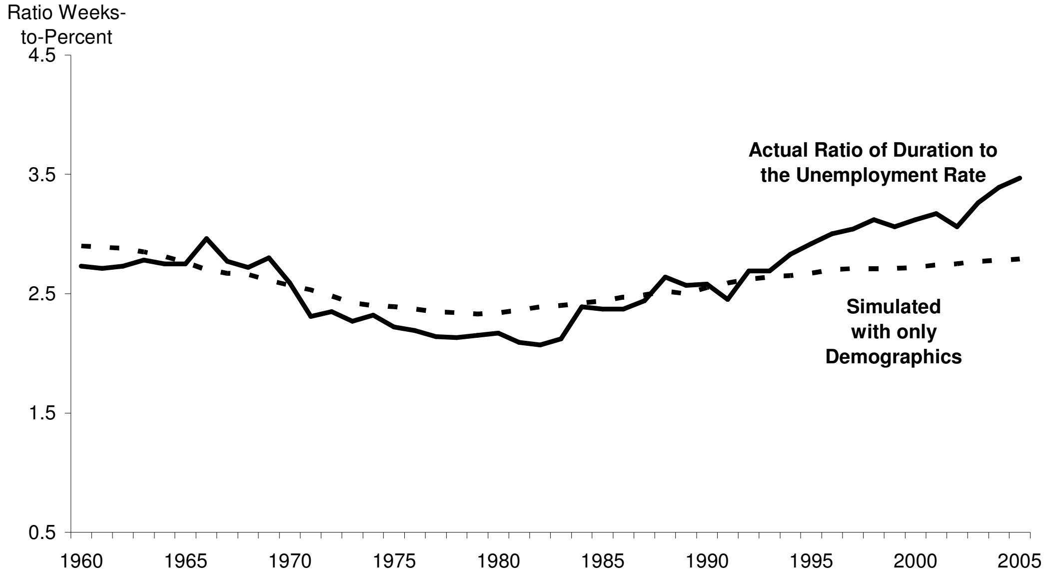
Sources: Bureau of Labor Statistics and authors' calculations.

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



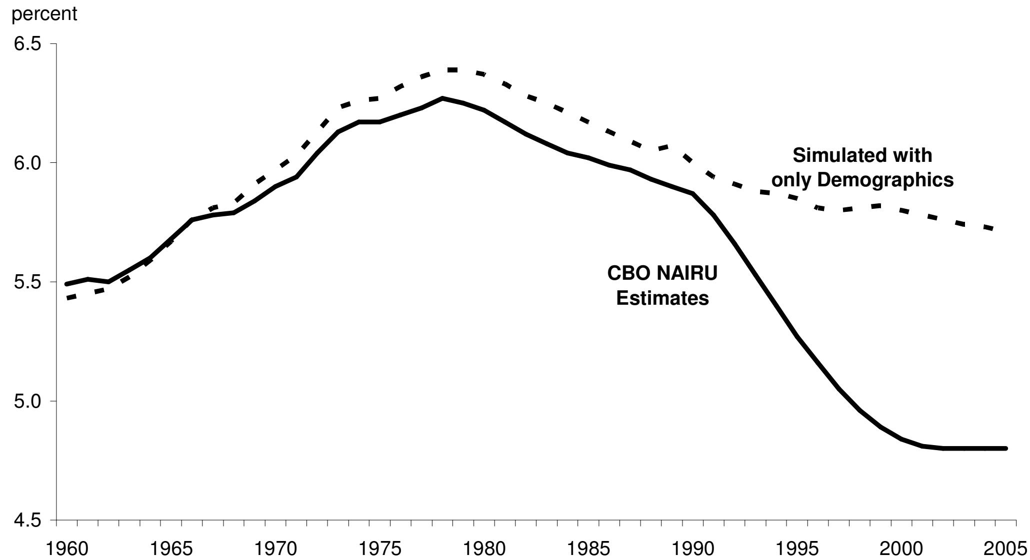
Source: International Survey Research .

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' calculations.

Figure 5: Demographic Shifts Stop Tracking Actual and Cyclically-Adjusted Separation Rates Since the Early 1990s

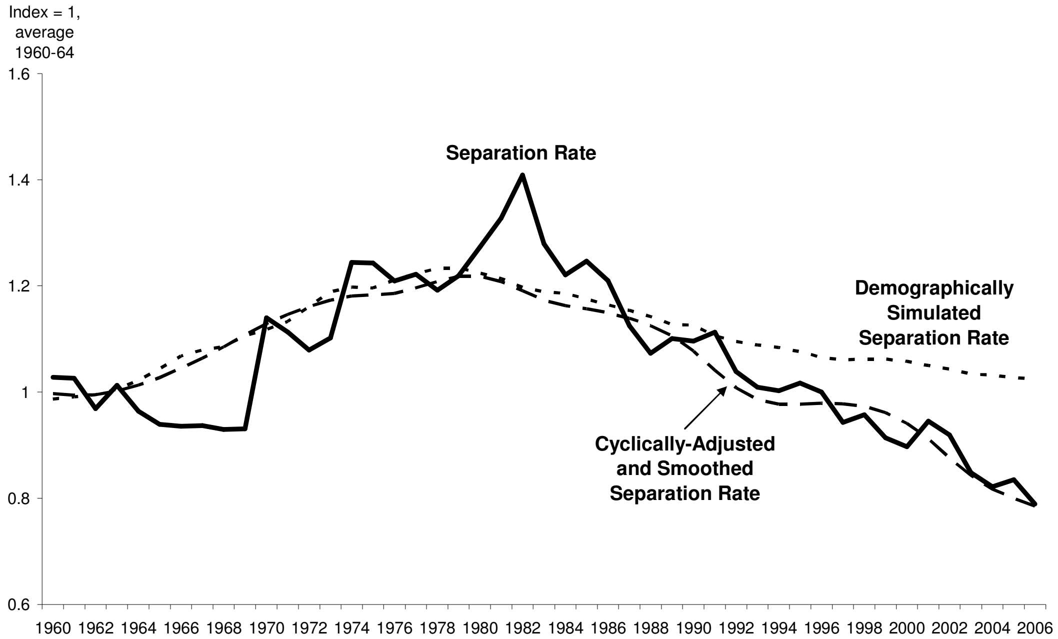
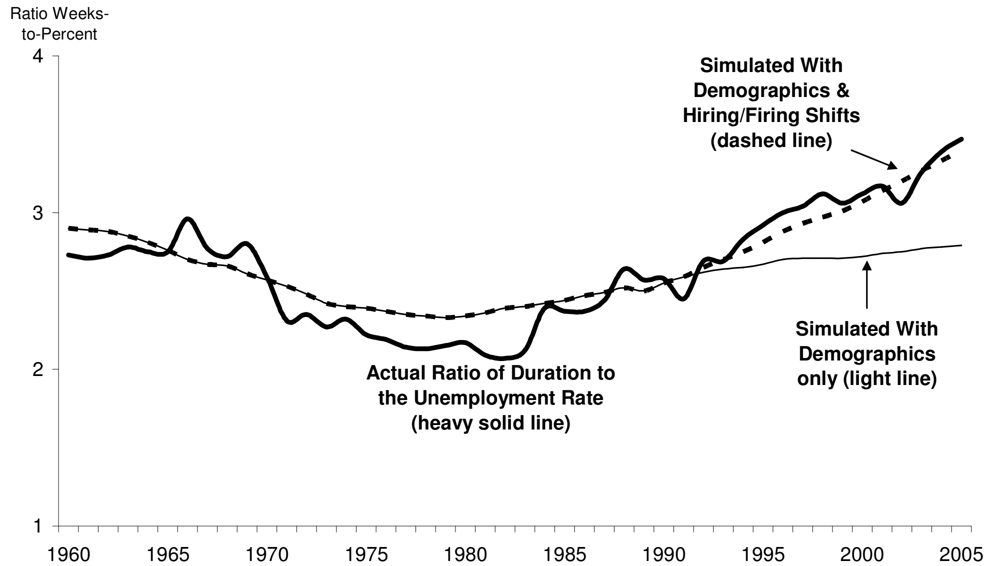
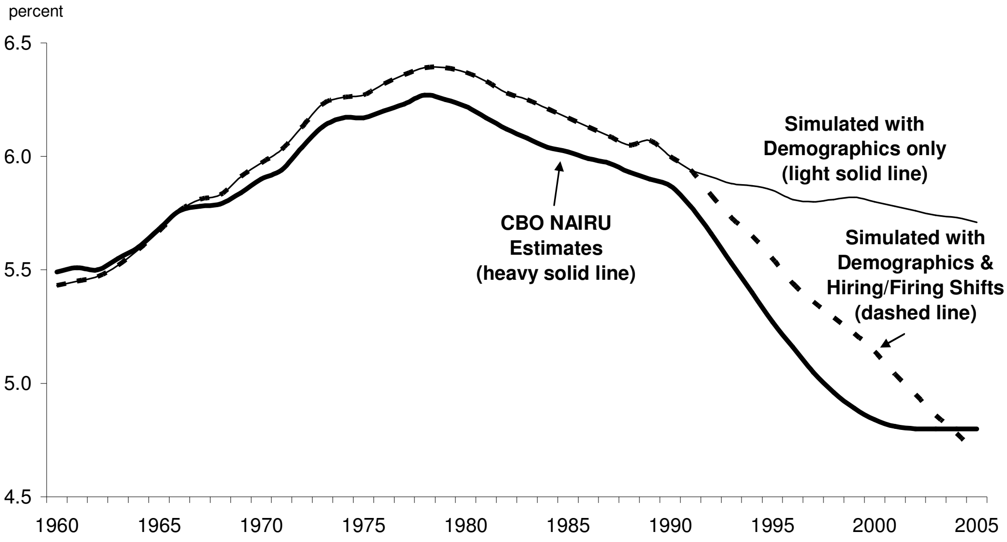


Figure 6: Demographics and Shifts in Turnover Behavior Track the mid-1990s' Relative Rise of Duration to the Unemployment Rate



Sources: Bureau of Labor Statistics and authors' calculations.

Figure 7: Demographics and Shifts in Turnover Behavior Help Track CBO NAIRU Estimates in the 1990s



Sources: CBO and authors' calculations.

Figure 8: In Core PCE Models, the NAIRU Varies with Duration

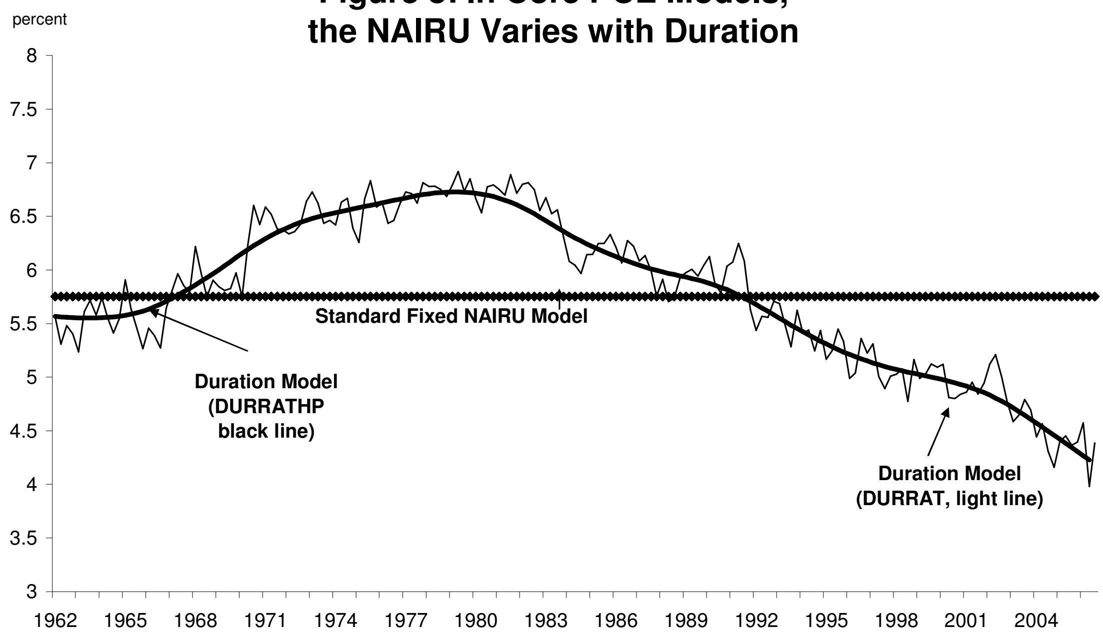


Table 1: Core PCE Inflation Regressions with Real Ex. Rate, Sample: 1962:Q2-2006:Q4[1]

	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**	0.0087**	0.0105**	0.0084**	0.0084**	0.0084**	0.0040**	0.0071*
	(5.01)	(5.02)	(3.59)	(2.74)	(4.74)	(2.66)	(5.64)	(4.64)
U_{t-1}	-0.0006**	-0.0008**	-0.0008**	-0.0008**	-0.0008**	-0.0008**	-0.0007**	-0.0008**
	(5.23)	(-6.34)	(-5.53)	(-5.50)	(-6.16)	(-5.34)	(-5.84)	(-6.31)
$NIXON_t$	-0.0038**	-0.0045**	-0.0038**	-0.0045**	-0.0042**	-0.0042**	-0.0039**	-0.0043**
	(-2.86)	(-3.41)	(-2.87)	(-3.36)	(-3.20)	(-3.15)	(-3.00)	(-3.32)
$NIXOFF_t$	0.0108**	0.0104**	0.0105**	0.0104**	0.0103**	0.0103**	0.0107**	0.0104**
	(7.25)	(7.15)	(7.16)	(7.13)	(7.08)	(7.06)	(7.32)	(7.13)
$ENERGY_t$	0.0002	0.0029*	0.0022	0.0029*	0.0028*	0.0028+	0.0020	0.0026+
	(1.26)	(2.08)	(1.61)	(2.07)	(1.99)	(1.97)	(1.49)	(1.91)
$\Delta RER8Q_{t-2}$	-0.0048*	-0.0050**	-0.0041*	-0.0050*	-0.0048*	-0.0048*	-0.0049*	-0.0051**
	(-2.43)	(-2.61)	(-2.08)	(-2.54)	(-2.50)	(-2.41)	(-2.54)	(-2.66)
$DURRAT_{t-1}$		-0.0015**		-0.0015*				-0.0009*
		(-3.39)		(-2.19)				(-2.29)
$DURRATHP_{t-1}$					-0.0014**	-0.0014+		
					(-3.12)	(-1.75)		
$AGE35+_{t-1}$			-0.00010*	0.000001		-0.00001		
			(-2.55)	(0.13)		(-0.01)		
NAIRU, 06:Q4	5.75%	4.39%	4.87%	4.37%	4.23%	4.22%	6.15%#	5.27%#
$\%\Delta U^{\bullet} / \%\Delta R$		0.810			0.792			
LM(1)	0.54	1.04	0.83	1.05	1.17	1.17	1.07	1.17
LM(2)	0.72	1.11	0.85	1.13	1.17	1.17	1.07	1.23
q(24)	18.23	15.82	14.82	15.93	16.28	16.26	17.12	17.08
R^2	.8882	.8950	.8919	.8943	.8939	.8932	.8920	.8948

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 2: Nominal Wage Inflation Regressions With Real Ex. Rate (NonFarm Business Sector), Sample: 1962:Q2-2006: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	0.0118**	0.0260**	0.0353**	0.0267**	0.0265**	0.0254**	0.0141**	0.0221**
	(5.91)	(7.84)	(5.76)	(4.04)	(7.81)	(3.73)	(7.06)	(7.58)
U_{t-1}	-0.0015**	-0.0021**	-0.0023**	-0.0021**	-0.0021**	-0.0021**	-0.0018**	-0.0020**
	(-5.75)	(-7.82)	(-7.23)	(-6.87)	(-7.84)	(-6.72)	(-6.94)	(-7.83)
$NIXON_t$	-0.0046+	-0.0062*	-0.0040	-0.0061*	-0.0055*	-0.0056*	-0.0048+	-0.0059*
	(-1.67)	(-2.40)	(-1.53)	(-2.30)	(-2.17)	(-2.16)	(-1.83)	(-2.31)
$NIXOFF_t$	0.0067*	0.0057*	0.0064*	0.0057*	0.0056+	0.0056+	0.0067*	0.0056+
	(2.18)	(1.99)	(2.16)	(1.99)	(1.95)	(1.93)	(2.25)	(1.96)
$ENERGY_t$	-0.0032	-0.0004	-0.0014	-0.0004	-0.0004	-0.0004	-0.0026	-0.0012
	(-1.10)	(-0.16)	(-0.49)	(-0.15)	(-0.14)	(-0.14)	(-0.91)	(-0.43)
$PROD12_{t-1}$	0.3361*	0.5282**	0.6501**	0.5373**	0.5372**	0.5228**	0.3352**	0.5106**
	(2.46)	(4.00)	(4.28)	(3.52)	(4.05)	(3.40)	(2.61)	(3.85)
$\Delta RER8Q_{t-2}$	-0.0045	-0.0047	-0.0019	-0.0046	-0.0042	-0.0044	-0.0049	-0.0048
	(-1.08)	(-1.22)	(-0.47)	(-1.14)	(-1.09)	(-1.10)	(-1.21)	(-1.25)
$PROFSHAR_t$	0.8759**	0.9046**	0.8701**	0.9034**	0.8974**	0.8990**	0.8717**	0.9009**
	(13.44)	(14.89)	(13.97)	(14.65)	(14.80)	(14.64)	(13.94)	(14.82)
$DURRAT_{t-1}$		-0.0044**		-0.0043*				-0.0029**
		(-5.17)		(-3.06)				(-3.65)
$DURRATHP_{t-1}$					-0.0046**	-0.0048**		
					(-5.18)	(-3.08)		
$AGE35+_{t-1}$			-0.0004**	-0.00002		0.00003		
			(-4.04)	(-0.12)		(0.19)		
NAIRU,06:Q4	-1.57%	4.27%	4.67%	4.30%	4.01%	3.99%	6.56%[#]	4.67%[#]
$\%\Delta U^{\bullet} / \%\Delta R$		-1.000			-1.024			
D.W.	1.62	1.87	1.78	1.87	1.87	1.83	1.74	1.87
LM(1)	7.00**	0.78	2.22	0.79	0.83	0.84	3.31+	0.79
LM(2)	9.72**	0.81	2.71	0.82	0.93	0.95	4.46	0.86
q(24)	42.11*	27.74	26.64	27.77	27.24	27.17	29.42	27.62
R^2	.7374	.7737	.7603	.7723	.7739	.7725	.7565	.7739

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] Other versions of efficiency wage theory are the gift-exchange model of Akerlof (1982, 1984) and the adverse selection model of Weiss (1980). In these models, efficiency is probably less dependent on the unemployment rate or the hiring rate than in the shirking and labor turnover models.

[2] As in Shimer (2007), it is assumed that workers transition between employment and unemployment and do not enter or leave the labor force. Shimer explores the implications of relaxing this assumption and finds that, "although this changes the level of job finding and exit probabilities, it does not quantitatively affect their fluctuations" (p. 2). In addition, Blanchard and Diamond (1990) show that approximately 50% of hires are from individuals classified as "not in the labor force," suggesting that the distinction between being unemployed and being out of the labor force is not important for many individuals.

[3] These calculations assume that the economy is in a steady-state equilibrium in which employment is constant.

[4] This appendix is available from the authors upon request. In the efficiency wage model, the elasticity of the NAIRU with respect to the separation rate is 0.943 (given the baseline parameters). In the matching model, this elasticity lies between 0.987 and 0.976 when reasonable values are chosen for the model's parameters.

[5] Estimates of the natural rate are available at www.cbo.gov/Spreadsheet/8565_table2-2.xls.

[6] These data were constructed by Robert Shimer. For additional details, please see Shimer (2007) and his webpage http://robert.shimer.googlepages.com/flows. The separation data are available at http://robert.shimer.googlepages.com/sep-prob.dat.

[7] Data on hiring probabilities are available at http://robert.shimer.googlepages.com/find-prob.dat.

[8] Figures reporting the results of these simulations are available from the authors upon request. In these simulations, there is one value of q for the entire economy, and this value of q depends on the percentage of workers above and below 35, as well as on the percentage in each demographic subcategory.

[9] *NIXON* equals 1 during the first two quarters of price controls (1971:3-71:4) and equals 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.

[10] 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.

[11] 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 President*, 1997). A productivity variable (*PROD*12) 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.

[12] Owing to this scaling, the reported coefficient differs to a large multiplicative extent compared to an earlier version of this paper when *PROD*12 was defined in percentage points and was an annualized rate.

[13] 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.

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

[15] 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."

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

[17] As discussed in Pissarides (2000), z includes unemployment benefits, income earned from temporary jobs when unemployed, the value of home production, and the value of leisure.

[18] Vacancy data were obtained from Figure 5 of Katz and Krueger (1999), which plots vacancy rates for years between 1960 and 1998. Years in this time frame in which the natural rate was within 0.5% of the CBO's estimate of the natural rate were 1960, 1962, 1963, 1964, 1971, 1972, 1978, 1979, 1987, 1988, 1990, 1995, 1996, and 1997. The average vacancy-unemployment ratio was approximately 0.3 both when the sample consists of all of these years and when the sample consists of the years in the 1960's.