# Declining labor force participation and its implications for unemployment and employment growth

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## Introduction and summary

The labor force participation (LFP) rate—the share of the working-age population that is either employed or jobless and actively looking for employment—has fallen from 66 percent at the beginning of the Great Recession in December 2007 to 62.7 percent in September 2014.[1] To some, this decline suggests the possibility that there may be labor market slack over and above that captured by the unemployment rate. The existence of such extra slack might imply that it would be appropriate for monetary policy to remain highly accommodative for longer than would otherwise be the case. However, to properly judge the extent to which the drop in the LFP rate reflects additional slack, one must account for the effects of several long-running trends not associated with the latest recession. Such pre-recession trends include the movement of baby boomers into retirement ages, long-running declines in the labor force participation of males of prime working age (25–54), the flattening out of once-rising female participation, sharp declines in teen participation, and the increasing participation of adults aged 55 and older. All but the last trend imply that a decline in aggregate LFP was to be expected even before the Great Recession began. Indeed, after rising from the 1960s through the 1990s, LFP has been falling since 2000, reflecting most of these factors.

In this article, we extend the methodologies of Aaronson and Sullivan (2001), Sullivan (2007), Aaronson, Davis, and Hu (2012), and Aaronson and Brave (2013) to provide estimates of the long-run trend rate of LFP[2] based on pre-recession data (data before 2008). Our models (with different specifications) suggest that the actual LFP rate as of the third quarter of 2014 is 0.2 to 1.2 percentage points lower than what would have been expected before the recession started, with our preferred model estimating the gap at the high end of this range.[3] We also provide a prediction of the LFP rate that would have been expected given the high unemployment rates of recent years and find that the actual LFP rate as of late is 0 to 0.8 percentage points lower than that benchmark, with our preferred estimate again being at the high end of the range. The results from our models suggest that there may indeed be greater slack in the labor market than is signaled by the unemployment rate.

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Our analysis is based on the full set of micro-level data on labor force participation collected in the U.S. Bureau of Labor Statistics' (BLS) *Current Population Survey* (CPS)—often referred to as the household survey—since 1982. These BLS data allow us to estimate statistical models that independently account for the long-running patterns we have mentioned. In particular, microdata allow our statistical models to identify life-cycle work patterns by very fine age groups and, further, to account for how specific cohorts follow these life-cycle patterns to varying degrees depending on when they were born. This is useful because cohorts that have high LFP early in their working careers tend to continue to have high LFP later in their careers as well.[4]

There have been important changes over time in these birth-cohort-specific LFP tendencies. On the one hand, successive cohorts of men, especially those with low levels of education, have had lower and lower LFP tendencies. On the other hand, for several decades successive cohorts of women tended to work more than earlier cohorts. However, those born after roughly 1960 did not show much further increase in LFP and the latest cohorts of women may even be showing some declines relative to earlier cohorts—similar to the pattern that has prevailed among men for several decades. Thus, as the women born before 1960 have exited their prime working years, their upward influence on women's LFP has largely disappeared.

Our models also allow LFP to vary by education level, reflecting the well-known positive association between educational attainment and LFP. During the latter half of the twentieth century, substantial increases in educational attainment were a factor in the long-running increase in LFP. More recently, however, educational improvement has slowed considerably, implying there is less impetus for LFP to rise. Additionally, we control for other factors that might drive participation decisions, including longer life spans, changes over time in the prevalence of young children (those under five years old), and factors such as the real minimum wage and the adult-to-teen wage ratio that might influence teen participation.

Finally, we allow LFP to vary with business cycle conditions. LFP tends to be below trend when unemployment is high—as has been the case for several years—and above trend when unemployment is low. Like Erceg and Levin (2013), we see evidence that this association is present with long lags. We also find that the strength of the relationship between unemployment and LFP varies with age, sex, and education levels. We exploit state variation in unemployment rates to estimate models that allow for long lags and demographic variation in the association between LFP and the unemployment rate. Properly controlling for unemployment has the important effect of stabilizing our estimates of the trend LFP rate. That is, we get almost the same trend LFP rate whether we end the estimation in 2007 or 2014 and very similar estimates even if we stop the estimation as early as 2002. Our findings contrast with some other estimates of the long-run trend rate of LFP, such as those provided by the BLS, which have changed considerably over time.

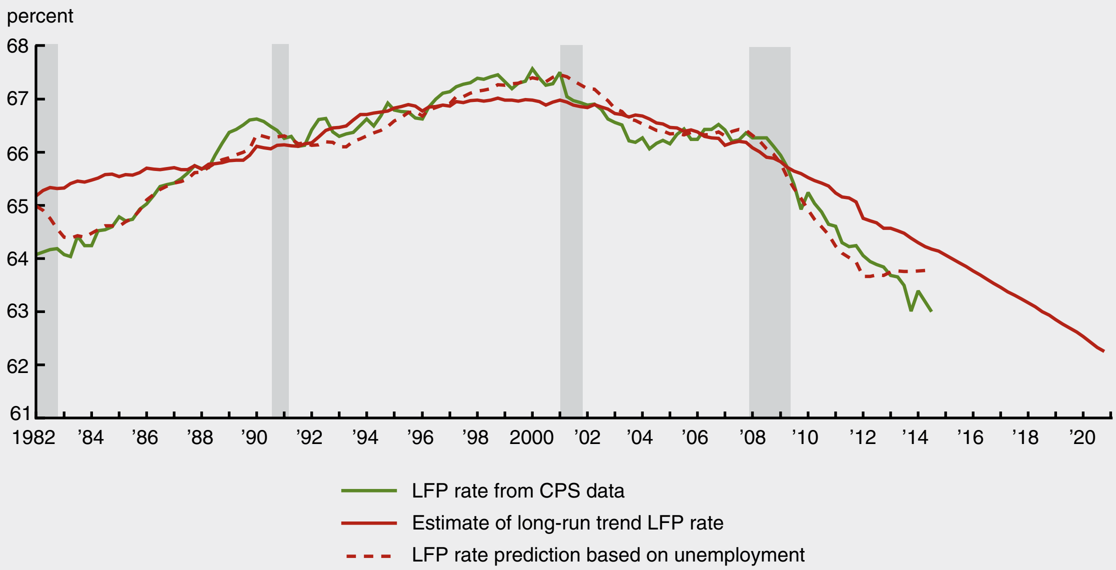
That said, our estimates of the gap between the actual and trend LFP rates depend on several particular modeling assumptions. So it is worth emphasizing that tweaks to the model naturally cause the results to vary somewhat. Most notably, our preferred model, whose results we highlight throughout this article, allows for separate birth-cohort coefficients for four age categories (16–24, 25–54, 55–79, and 80 and older[5]). This model, which we call our "baseline," implies that the actual LFP rate as of the third quarter of 2014 is about 1.2 percentage points lower than the trend LFP rate. If, instead, we force the cohort coefficients in the model to be the same for ages 16–79 (as in the "pooled model"), the LFP gap falls to around 0.2 percentage points. We discuss these different models and other robustness checks in detail later in the article.

Figure 1 shows the history of the actual LFP rate data from the CPS (solid green line), along with our baseline estimate of the long-run trend LFP rate (solid red line) and the corresponding prediction of the LFP rate given the recent history of state-level unemployment gaps[6] (dashed red line). According to our estimates, after rising for many years, the trend LFP rate began to decline after 2000. Recently, according to our baseline model, that decline has accelerated to about 0.3 percentage points per year—an annual rate of decline that our model suggests will persist for the foreseeable future. By 2020, our baseline model predicts the trend LFP rate to be 62.3 percent, its lowest level since the mid-1970s. The 2020 rate is even lower when we force the cohort coefficients in the model to be the same for ages 16–79 (as in our pooled model, whose results are not shown in figure 1).

As of the third quarter of 2014, our baseline estimate of the trend LFP rate is 64.2 percent, about 2 percentage points below our estimate of this trend rate at the end of 2007. However, while the trend rate fell by about 2 percentage points, the actual LFP rate dropped even more, leaving it 1.2 percentage points below the long-term trend. Additionally, until the third quarter of 2013, the LFP rate had followed relatively closely its predicted path based on prevailing labor markets conditions. However, since then, the actual LFP rate has dipped below even the rate predicted

FIGURE 1

Labor force participation (LFP) rates, 1982–2020




Notes: The figure plots quarterly data for those aged 16 and older over the period 1982:Q1–2020:Q4. The long-run trend LFP rate is based on data through 2007:Q4. The LFP rate prediction based on unemployment is the rate based on contemporaneous state unemployment gaps (and their lags). See the text for further details. The shaded bars indicate recessions as defined by the National Bureau of Economic Research.

Source: Authors’ calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*.

with the high unemployment rates of the past several years (see figure 1). This gap suggests there is an extra margin of slack over and above what one would infer from the unemployment rate alone.

Our results also have implications for the natural rate of unemployment[7] that may suggest greater labor market slack. In particular, the decline in the trend LFP rate that we find has not been uniform across different populations. Certain groups, such as those under age 25, have seen particularly large drops in LFP, while the LFP of other groups, such as those over age 54, has actually increased. In addition to these uneven LFP trends, educational attainment, while not improving as rapidly as in earlier decades, has steadily advanced. These trends have led the labor force to be somewhat more heavily weighted toward groups that tend to have low unemployment, such as older people and those with higher levels of educational attainment. We estimate that on their own, these developments would have lowered the natural rate of unemployment by about 0.3 percentage points since 2007 and 0.6 percentage points since 2000. Recent estimates of the natural rate have focused on developments such as the increase in long-term unemployment that some argue have raised the natural rate. The demographic and educational effects on the natural rate we document here are large enough to offset most of those adverse influences, suggesting that the natural rate may be lower than is often assumed.

Another implication of our results is that once employment and output have returned to their long-run trends, they will grow more slowly than in the past. All else being equal, an LFP rate that is declining by 0.3 percentage points per year translates into 0.5 percentage points less growth in hours worked per year and thus, if productivity growth is unchanged, 0.5 percentage points less potential output growth per year, compared with an economy with a flat LFP rate. The slow fall in the natural rate of unemployment implied by our results offsets a small portion of those effects. In combination with the U.S. Census Bureau’s assumption about population growth, our results imply that trend payroll employment growth[8] will fall to under 50,000 jobs per month later in the current decade. However, that “normal” employment growth rate will only become apparent in the data after a still sizable employment gap (that is, the difference between the actual and trend level of total payroll employment) that opened up during 2008–09 is finally closed.

To understand why LFP has been running below expectations, it is helpful to identify the groups for which the LFP gap has been especially large. Much of the surprise has occurred among adults without a college degree (high school dropouts, in particular)—whose actual LFP rates have dropped by even more than our estimates of the trend rates. At the end of the article, we speculate on possible reasons for these discrepancies and whether they might be resolved eventually or instead turn out to be signs that the model might be missing important developments.

Finally, we should add a note of caution about these LFP forecasts. The statistical models underlying our estimates of the trend in LFP and other variables mainly just extrapolate long-running trends. We do not attempt to explain the decline in LFP at the level of the underlying supply and demand for labor. Thus, the trends we identify could be altered by policy changes in such areas as disability insurance or education policy. It is also possible that a continued drop in LFP might elicit endogenous macroeconomic responses—for instance, more rapid wage growth—that might limit the phenomenon in the future. Developing a deeper understanding of the drop in LFP might thus be a fruitful area for future research.

In the next section, we briefly explain the key reasons behind long-running trends in LFP over roughly the past 60 years. We then describe the data we use and outline the methodology behind our estimate of the trend LFP rate. Afterward, we present our results, beginning with our aggregate estimates and then moving on to decompositions that quantify the demographic (age and sex), "behavioral" (for example, educational attainment, fertility rate, and life expectancy), and business cycle factors driving our findings; we follow this up with a discussion on the robustness of our results. In the final sections, we examine the impact of our LFP results on the estimate of the natural rate of unemployment and describe our estimates of trend payroll employment growth.

## Background

The LFP rate began to steadily increase in the mid-1960s, persistently expanding through the 1990s and peaking at 67.3 percent in 2000. According to the BLS, as of September 2014, however, the LFP rate is 62.7 percent—back toward the levels that were prevalent in the late 1970s.[9]

Many factors can be associated with the upsurge in LFP from the mid-1960s through the late 1990s and its subsequent drift back down since 2000. Perhaps the upswing and certainly the more recent downward pattern mirror the life-cycle work decisions of the large baby boom cohort, born during the two decades following World War II. Like every birth cohort, the LFP of baby boomers follows a distinct lifetime pattern. Labor force participation is low for teenagers, rises as individuals finish school in their late teens and early twenties, flattens out for those in their prime working years when work decisions are particularly insensitive to wages and economic conditions, and then falls for those in their late fifties and sixties as they enter retirement (see the blue and red lines in figure 2).[10] The baby boomers entered their prime working years during the 1970s and 1980s; and because of their sheer numbers, they caused an upsurge in aggregate LFP that lasted for decades. However, starting around 2000, a growing number of baby boomers reached their fifties and started to transition out of the labor force. Today, those same workers are now in their sixties and seventies, when LFP is much lower. To help make this point, we feature orange bars in figure 2 representing changes in the share of the working-age population for different age groupings over the years 2010–15.

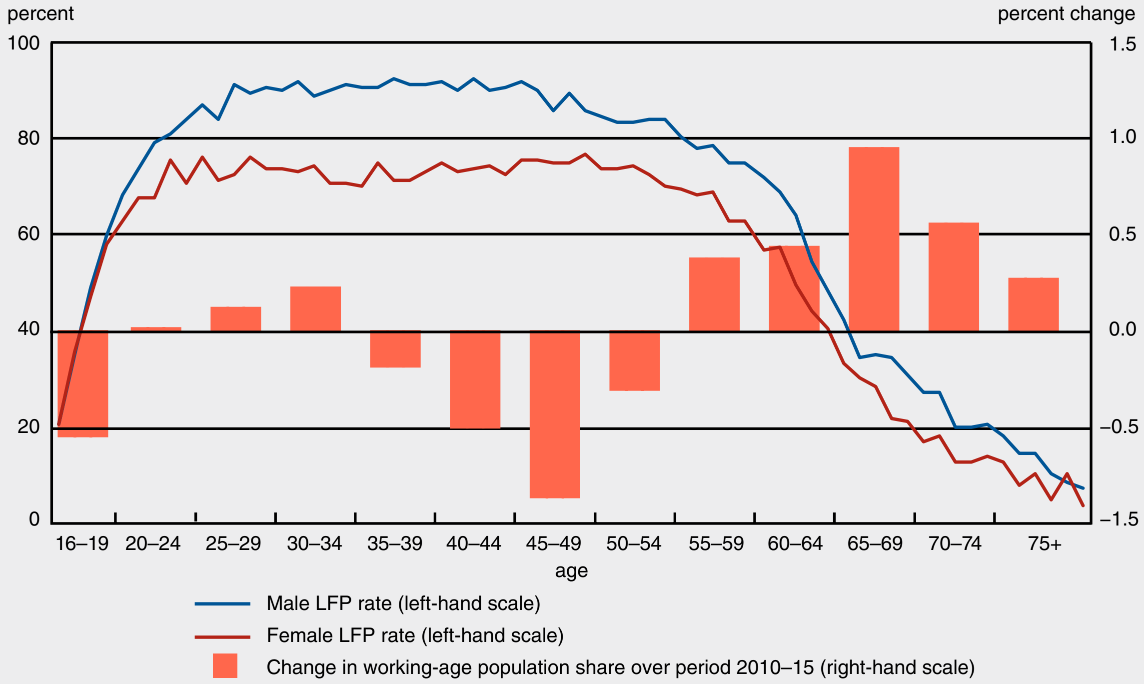
To quantify the importance of population aging, we compare in figure 3 the actual aggregate LFP rate with the LFP rate implied by demographic change—specifically, one that holds age-sex groups' actual LFP rates fixed at 2007 levels while allowing their population shares to vary according to the actual data and U.S. Census population projections. Since 2007, the actual aggregate LFP rate fell by 3.2 percentage points, while the rate implied by demographic change fell by 1.8 percentage points. The difference means that changing demographics alone explain only about half of the decline in LFP since 2007.

As figure 3 makes clear, even *within* demographic groups, there have been important changes in labor force attachment over time. The dramatic increase in the number of working women (red line in figure 4) was clearly a driving force behind rising LFP rates during the second half of the twentieth century. Only one in three women were in the labor force in 1948, but by the late 1990s, the female LFP rate was roughly 60 percent. However, by the end of the twentieth century, the female LFP rate had, more or less, leveled off. The female LFP rate has even declined some since the onset of the latest recession.[11]

By contrast, the male LFP rate has been on an uninterrupted decline since the 1950s, falling from 86.7 percent in 1948 to 69.2 percent in 2014 (blue line in figure 4). There is significant uncertainty about the precise cause of this secular decline, but researchers have linked it to stagnating overall real wage increases or declining real wages for low-skill workers; changes in safety net programs, in particular Social Security Disability Insurance (DI) and Supplemental Security Income (SSI); and increases in the labor force participation of women.[12] Over the past decade, the disappearance of manufacturing and other "middle-skill,"

**FIGURE 2**

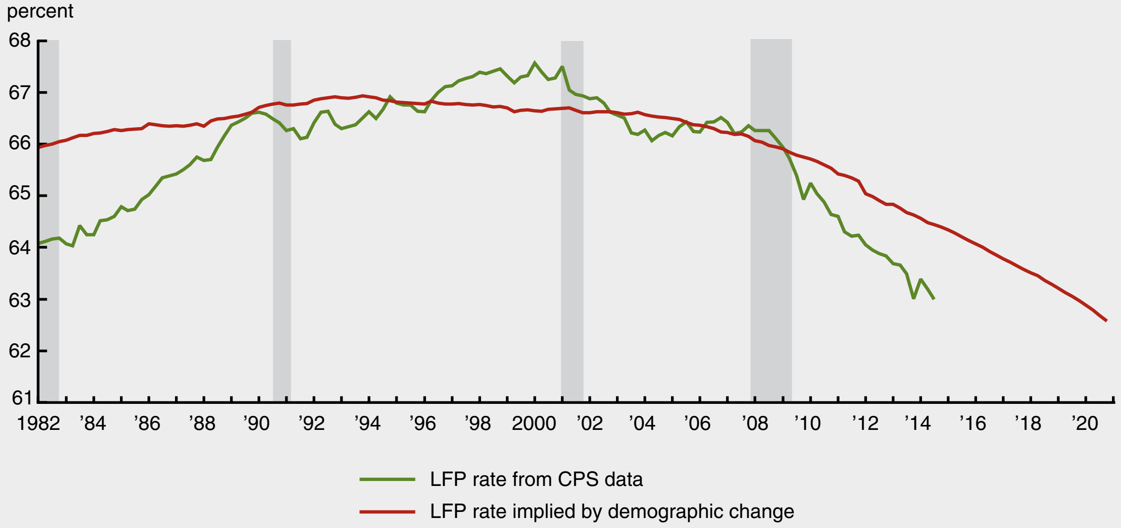
**Labor force participation (LFP) rates, by age and sex, and working-age population change, by age**




Note: The LFP rates for males and females aged 16 and older (of working age) are for 2014:Q3.
Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*.

**FIGURE 3**

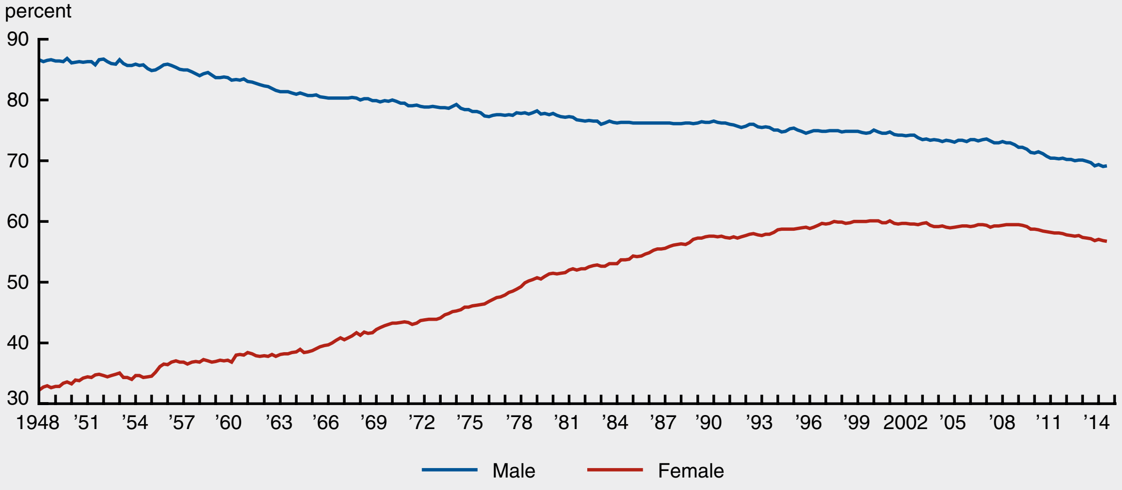
**Labor force participation (LFP) rate and LFP rate implied by demographic change, 1982–2020**




Notes: The figure plots quarterly data for those aged 16 and older over the period 1982:Q1–2020:Q4. The LFP rate implied by demographic change is the LFP rate that holds the actual LFP rates fixed at their 2007 levels for single age groups by sex while allowing group-specific population shares to vary according to the actual data and U.S. Census population projections. The shaded bars indicate recessions as defined by the National Bureau of Economic Research.
Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey* (CPS).

FIGURE 4

Labor force participation rate, by sex, 1948–2014




Note: The figure plots official quarterly data for those aged 16 and older over the period 1948:Q1–2014:Q3.
Source: U.S. Bureau of Labor Statistics, *Current Population Survey*, from Haver Analytics.

middle-income jobs may have contributed.[13] Indeed, the most recent recession and slow recovery have been particularly difficult for men; the LFP rate for men has fallen by 4.0 percentage points since December 2007—1.3 percentage points more than for women.[14]

Moreover, participation for many narrow age-sex groups has been changing over time (see panels A through F of figure 5). Because prime-working-age participation rates (shown in panels C and D of figure 5) echo the aggregate gender-specific trends (down for men, but up and then flat or down somewhat for women), we will focus on trends for teens and older individuals.

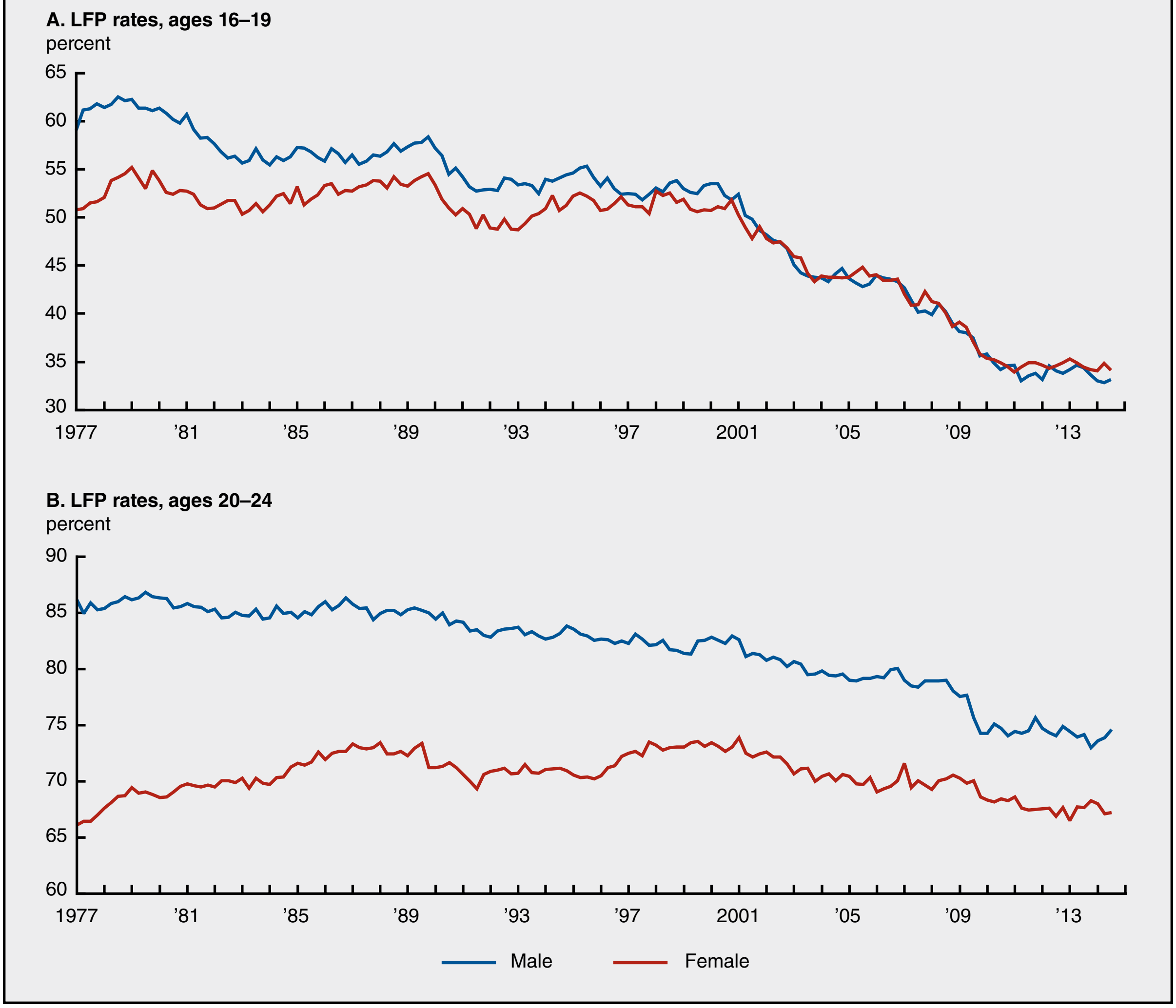
Teen participation has declined dramatically since the late 1970s (panel A of figure 5), particularly during the past decade (Aaronson, Park, and Sullivan, 2006). One explanation is that teens are spending more time in school, especially during downturns when the opportunity cost of schooling is low (Barrow and Davis, 2012). In addition, Smith (2011) argues that the decline of low-skill jobs and middle-skill, middle-income jobs has pushed workers who used to fill those positions into other jobs that have traditionally been performed by teens. Increased immigration of low-skill workers could have the same impact on teen jobs (Smith, 2012).

At the other end of the working life (panels E and F of figure 5, p. 108), retirement is starting at later ages. For example, relative to the first quarter of 2000, an additional 6.0 percentage points of men aged 60–64 and 10.6 percentage points of women aged 60–64 are working.[15] Research suggests several factors have contributed to people working longer. Improvements in health technology may have boosted labor force participation directly, by improving the health and longevity of the work force, and indirectly, by requiring individuals to work longer to accumulate the wealth to support lengthier retirements. Changes to private pensions and Social Security (Blau and Goodstein, 2010; and French and Jones, 2012), as well as volatile retirement account balances and housing prices (French and Benson, 2011), may have increased the need to postpone retirement, particularly early in the expansion following the Great Recession, when household wealth dipped, a pattern that may be reversing as net worth recovers (Fujita, 2014). Note also that the increase in LFP has been particularly strong among older women. That might be a consequence of the rising participation of women in the late twentieth century; cohorts that worked more throughout their prime working years carry that work behavior forward into older ages. Accommodating the observed demand for elongating careers, part-time "bridge" jobs have become a more common way to transition slowly into retirement (Ruhm, 1990; Schirle, 2008; and Casanova, 2013).

A final factor in our analysis that is not considered in other LFP studies (for example, BLS studies such as Toossi, 2004, 2005, 2007; Aaronson et al., 2006; and Aaronson et al., 2014) is education. Rising rates of return to skills during the 1980s and 1990s encouraged human capital investment (Katz and Murphy, 1992; and Katz and Autor, 1999), resulting in a shift

**FIGURE 5**

**Labor force participation (LFP) rates, by age and sex, 1977–2014**




away from occupations that tend to have shorter average career lengths. During the prime working years, the relationship between educational attainment and work is unambiguously positive (see figure 6). For instance, at age 40, the LFP rates for male high school dropouts, high school graduates, and college graduates are 84.7 percent, 87.8 percent, and 95.1 percent, respectively (see figure 6, panel A). Moreover, individuals with less education tend to retire earlier. At age 62, the LFP rate for male high school dropouts, high school graduates, and college graduates is 53.4 percent, 58.4 percent, and 73.8 percent, respectively. Any feature of the labor market that encourages human capital investment will likely result in higher aggregate labor force participation down the road.

## Methodology

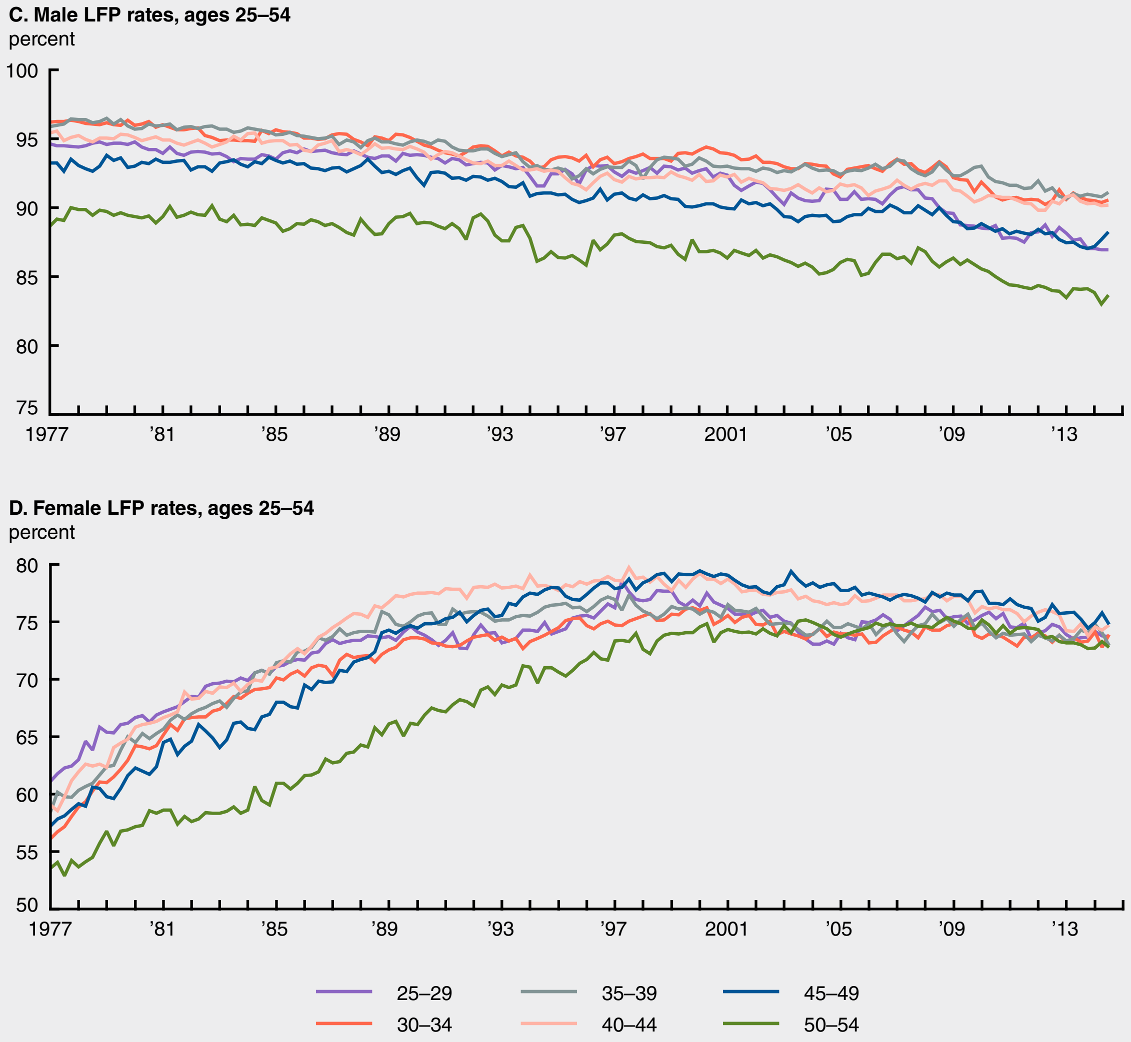
To measure the trend LFP rate, we estimate a statistical model of LFP that is capable of simultaneously considering various explanations based on demographic (age and sex), "behavioral" (for example, educational attainment, fertility rate, and life expectancy), and business cycle factors. First, we describe the data and then the statistical methodology.

### *Data*

Our LFP estimates are derived from the basic files of the U.S. Bureau of Labor Statistics' *Current Population Survey*. The CPS—the source of such well-known statistics as the unemployment rate and the labor force participation rate from the BLS—is a monthly, nationally representative survey of approximately

**FIGURE 5 (CONTINUED)**

**Labor force participation (LFP) rates, by age and sex, 1977–2014**




60,000 households conducted by the U.S. Census Bureau. Participating households are surveyed for four consecutive months, ignored for the next eight months, and then surveyed again for four more straight months. Important for our purposes is the fact that basic demographic data, such as age, sex, and race, as well as educational level and labor market status, are collected in the CPS. In the subsequent analysis, we use the microdata from January 1982 through September 2014.[16]

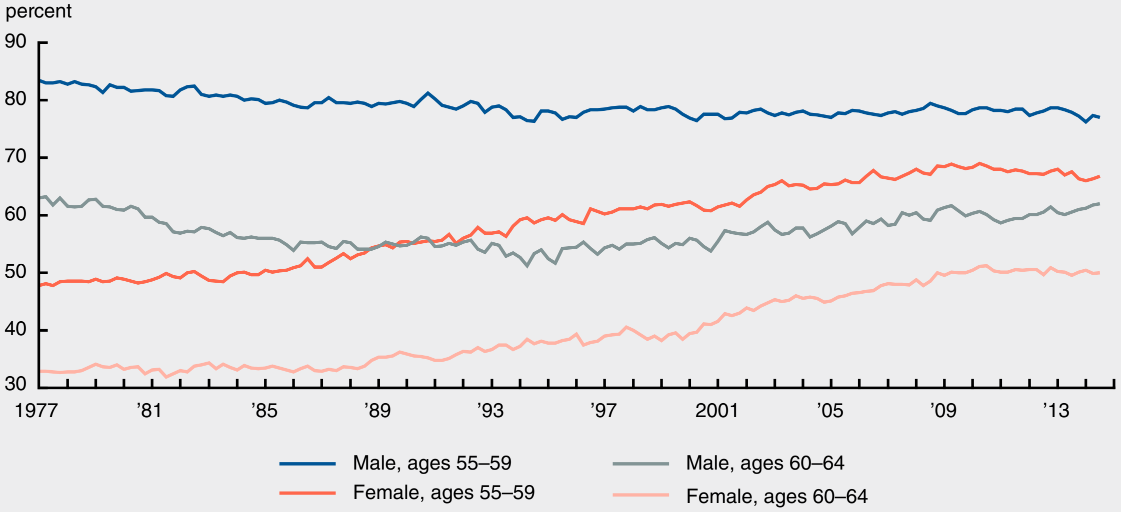
While the CPS contains information on many of the key determinants of labor force participation, we supplement our analysis with additional data to create controls used in the estimation of our statistical models.[17] Our additional controls are as follows:

- *The natural rate of unemployment, or the nonaccelerating inflation rate of unemployment (NAIRU).* We use the Congressional Budget Office's (CBO) calculation of the short-run NAIRU.[18]
- *State unemployment rate.* Since the state unemployment rates tabulated from the CPS data can be quite noisy, especially for small states, we use the state-level unemployment statistics published by the BLS. The BLS series is estimated using the CPS but augmented with data on unemployment insurance claims and payroll employment counts.[19]

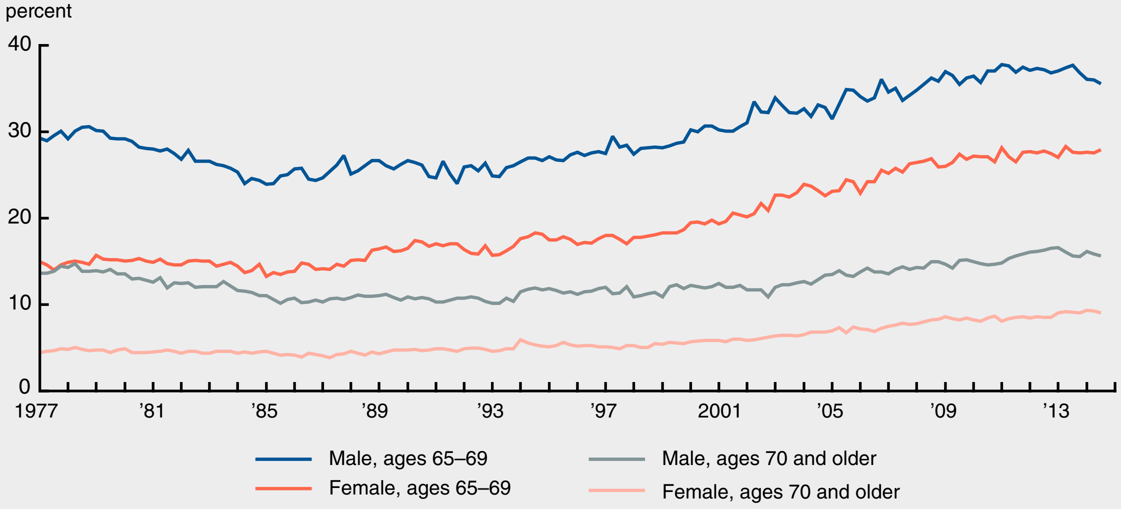
**FIGURE 5** (CONTINUED)

**Labor force participation (LFP) rates, by sex and age, 1977–2014**

**E. LFP rates, ages 55–64**




**F. LFP rates, ages 65 and older**




Note: Each panel plots official quarterly data over the period 1977:Q1–2014:Q3.
Source: U.S. Bureau of Labor Statistics, *Current Population Survey*, from Haver Analytics.

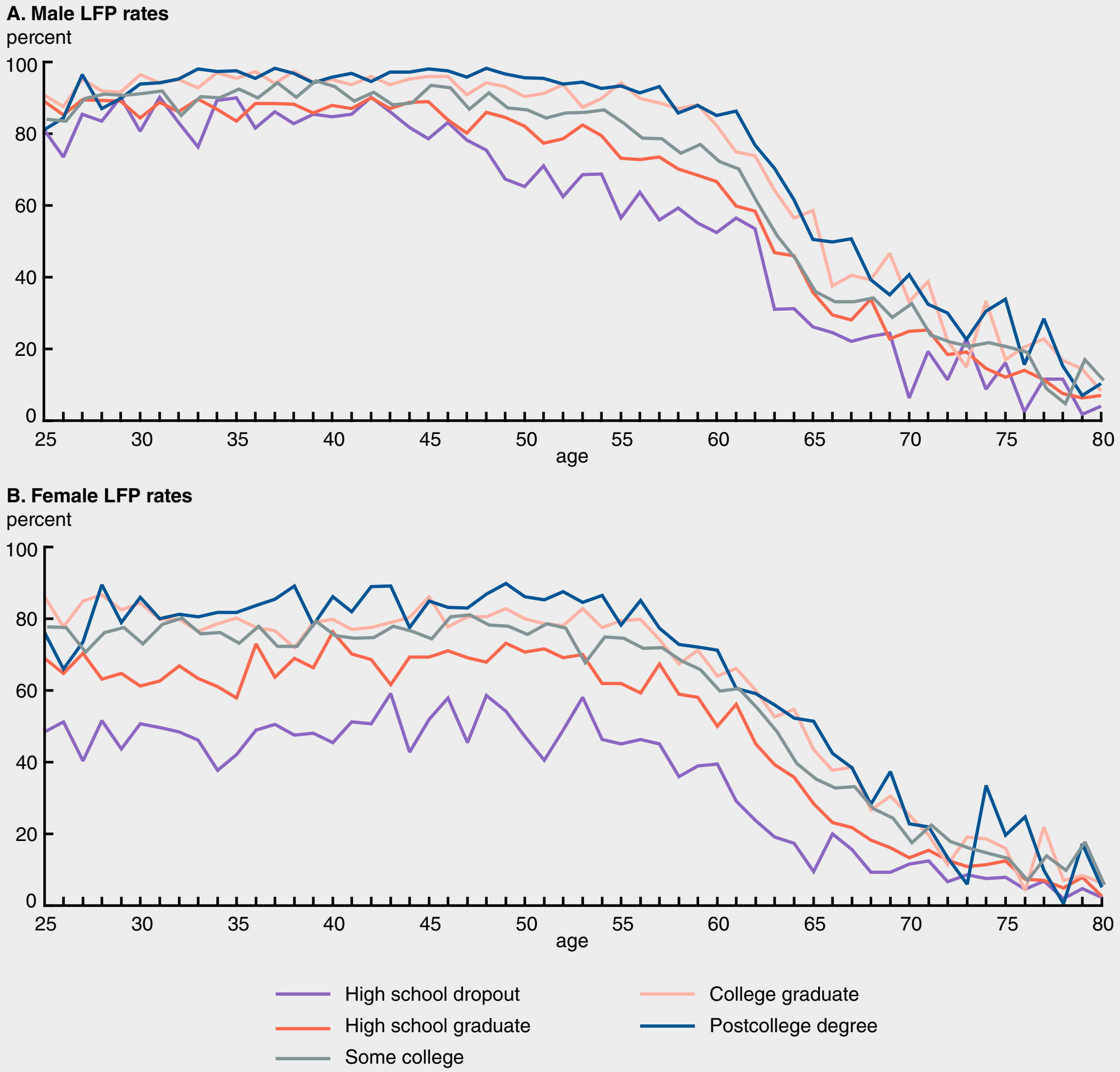
- *Minimum wage.* State minimum wage data are taken from the January issues of the BLS's *Monthly Labor Review* supplemented with minimum wage histories reported at the U.S. Department of Labor website.[20] The minimum wage data are deflated using the BLS's Consumer Price Index for All Urban Consumers (CPI-U).
- *Life expectancy.* Life expectancy, by sex and age, is taken from the Social Security Administration life tables (as in Bell and Miller, 2005) for the years 1980–2020. Missing years are linearly interpolated.

### *LFP model for ages 16–79*

Our baseline logistic regression model associates the probability that an individual aged 16–79 is in the labor force with that individual's sex, age, year of birth, race, and education level, as well as the economic

**FIGURE 6**

**Labor force participation (LFP) rates, by sex and education level, in 2014:Q3**




Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*.

conditions facing that individual and a few covariates specific to his or her age group:

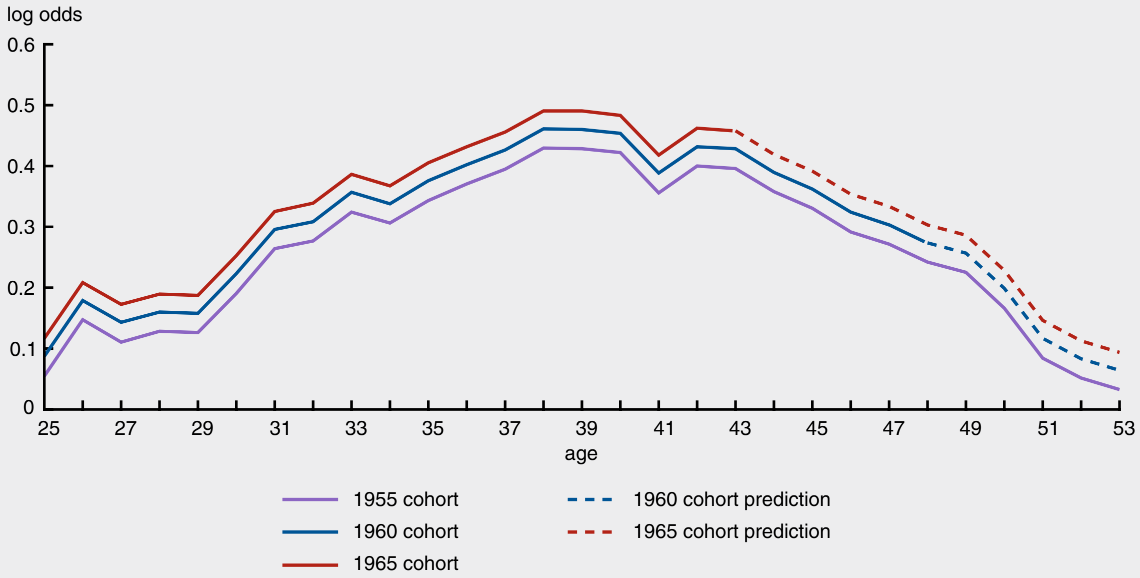
$$\log\left(\frac{p_{sebai}}{1-p_{sebai}}\right)=\alpha_{sea}+\beta_{seb}+w_{(a+b)i}\lambda_{se}+x_{sebai}\gamma_{se}+z_{seba}\delta_{se},$$

where $p_{sebai}$ is the probability individual $i$ of sex $s$ and education level $e$ born in year $b$ is in the labor force at age $a$. The left-hand side, $\log\left(\frac{p_{sebai}}{1-p_{sebai}}\right)$, is the log odds of being in the labor force.

The determinants of LFP (the right-hand side of the equation) include the key characteristics that drive the decision to work. For each sex-education-level group, a series of indicators for every single year of age, $\alpha_{sea}$, accounts for the typical lifetime pattern of labor force participation (see figure 6). A second series of indicators $\beta_{seb}$ are a full set of year-of-birth indicators for each sex-education-level group. According to the model, every birth cohort follows the same basic life-cycle pattern implied by $\alpha_{sea}$ but at a uniformly higher or lower level in terms of the log odds. This adjustment

FIGURE 7

Log odds of labor force participation (LFP) of unmarried white female high school dropouts aged 25–54 without a young child




Notes: The figure plots the log odds of LFP of unmarried white female high school dropouts aged 25–54 without a young child (under five years old) based on model estimates using quarterly data through 2007:Q4. The dashed lines represent projections.
Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*.

might reflect opportunities, preferences, and norms that are specific to particular birth cohorts.

To see the intuition underlying the cohort method for forecasting LFP, consider, for example, the problem of forecasting in 2007 the future participation of women who were then 42. We can compare the LFP rates at ages 25–42 of these women, who were born in 1965, against those of earlier birth cohorts at the same ages. The cohort method assumes that the average difference in LFP rates of these women and those of earlier cohorts will persist beyond age 42, allowing us to forecast their labor force participation for the remainder of their lives.

The idea is illustrated in figure 7 for one particular demographic group. This figure plots the predicted log odds $\left(\log\left(\frac{p_{sebai}}{1-p_{sebai}}\right)\right)$ of being in the labor force at ages 25–54 for unmarried white women without young children (under five years old) and without a high school diploma, born in 1955, 1960, and 1965, based on estimates using data through 2007. Note first that the age-to-age patterns for each line have a nearly identical shape. The only difference is that the lines are shifted up or down, with the size of that shift determined by our estimate of the cohort effect, $\beta_{seb}$. Through 2007, the cohort born in 1965 has been more likely to work at the same ages than the cohort born in 1960, which has been more likely to work than the cohort born in 1955, again at the same ages. By 2007, those in the 1965 cohort are only 42 years old. To forecast their LFP for the remainder of their careers, we assume the cohort difference up to age 42 will persist into older ages but the pattern over the remaining life cycle will look like that of past cohorts. This allows us to trace out their future participation (in a dashed line) for ages that we have yet to observe.

The cohort-based approach has an advantage over an alternative strategy (used most prominently by the BLS, as in Toossi [2004, 2005, 2007]) that bases the forecast on an extrapolation of the time series of the trend LFP rate for each age-sex group (using the last 13 years of data). A drawback of the BLS methodology is that it mixes different cohorts of women together, which could be problematic during periods when the level of trend LFP might be changing. That has been the case for much of the past few decades.

An individual's labor force participation decision will also be affected by the state of the economy and labor market conditions, $w_{(a+b)i}$. Statistical models of this sort (for example, Aaronson and Sullivan, 2001; and Aaronson et al., 2006) have typically relied on the contemporaneous gap between the actual unemployment rate and trend unemployment rate (that is, the natural rate of unemployment) to measure the state of the economy. However, this might misstate the role of the labor market for at least three reasons. First, Erceg and Levin (2013) provide evidence that LFP responds to unemployment with long lags. Consequently, we also account for the unemployment gap over the past three years—specifically, the average of this gap over the past zero to three quarters, four to seven quarters, and eight to 11 quarters. Like Erceg and Levin (2013), we find that there are substantial lags in the effects of unemployment rates on labor force participation. Second, labor market conditions vary substantially on a geographic basis, with some parts of the country experiencing more distress than others at a given date. Thus, our baseline model utilizes a state-level unemployment gap to account for more geographically detailed labor market conditions. Third, indicators other than the unemployment rate may be necessary to characterize the tightness of the labor market. Once we account for lagged state-level unemployment gaps, we find that the unemployment rate does an adequate job in characterizing labor market conditions. However, later on, we explore the robustness of our LFP results to the use of other measures, such as the national median length of unemployment.[21]

Finally, we introduce additional conditioning variables that are common to all demographic groups, $x_{sebai}$, and specific to certain age groups, $z_{seba}$. The main covariate common to all demographic groups is indicators for race. For 16–24 year olds, we also control for the real minimum wage and the hourly wage ratio of 16–24 year olds (youths) to 25–54 year olds (adults).[22] A higher minimum wage acts to encourage labor force participation, but perhaps to reduce the employment of teens (Neumark and Wascher, 2008, and references therein). Similarly, the overall ratio of teen to adult wages influences the market for teen employment (Aaronson, Park, and Sullivan, 2006; and Smith, 2011). For 25–54 year olds, we augment the model to include indicators for being married with a young child (under five years old), being married with no young child, and being unmarried with a young child (the omitted category is being unmarried with no young child). The impact of childbearing, particularly among women, can be seen in the dip in the LFP rate in the late twenties and early thirties for women (Leibowitz and Klerman, 1995; and Blau, 1998). Finally, we include measures of gender-specific life expectancy for 55–79 year olds to account for the delay in retirement caused by longer life expectancy.[23]

For flexibility, the baseline model is estimated separately by combinations of age, sex, and education level groups. Specifically, we break up the sample into 28 combinations of age (16–24, 25–54, and 55–79), sex, and educational attainment (high school dropout, high school graduate, some college, college graduate, and postcollege degree).[24] This allows the cohort effects and coefficients on other controls to flexibly vary across these groups. In particular, note that in the baseline, for each sex-education-level group, the parameters $\lambda_{se}$ and $\gamma_{se}$ also vary by the three age groups.[25] Later, we describe how the results change when we estimate the model forcing the cohort coefficients, $\beta_{seb}$, to be the same for the different age ranges.

### *LFP model for ages 80 and older*

A key feature of the baseline model for 16–79 year olds is the capacity to differentiate age and cohort effects. Unfortunately, this is not possible for individuals aged 80 and older because the CPS does not distinguish age beyond 80 in some years. Therefore, for those aged 80 and older, we replace age and cohort effects with a linear time trend:[26]

$$\log\left(\frac{p_{seti}}{1-p_{seti}}\right) = t\theta_{se} + w_{ti}\lambda_{se} + x_{seti}\gamma_{se} + z_{set}\delta_{se},$$

where $t$ indexes calendar time.[27]

The age group that is 80 and older is a very small share of the work force (about 4.5 percent of the working-age population and 0.43 percent of the employed in 2014).[28] Therefore, when we combine our LFP estimates of the 16–79 and 80-and-older populations, the resulting LFP rate for those aged 16 and older is not sensitive to the precise specification of the worker model for those aged 80 and older.

### *Estimation of the trend LFP rate*

The models are estimated using the CPS for the years 1982–2007. The additional exogenous variables are included at the quarterly frequency.[29] The trend rate is computed as the predicted LFP rate free of any variation due to the business cycle. In particular, we apply the coefficient estimates from the logit models we have just described to the data in order to predict age-, sex-, and education-level-specific group trend LFP rates $\hat{P}_{sea}$ assuming that the economy is at its current estimate of the natural rate of unemployment (that is, the current and lagged unemployment gaps $w_{(a+b)i} = 0$). The aggregate trend LFP rate is then the sum of the weighted group-specific trend LFP rates, where weights are allowed to vary over time based on a group's share of the overall population in that year. From the model, we can also compute a predicted LFP rate based on contemporaneous and lagged state unemployment gaps. This measure, which we label the "LFP rate prediction based on unemployment" (see figure 1, p. 102), reveals whether the actual LFP rate is unusually high or low given the present labor market situation as summarized by the history of unemployment gaps. In this way, it serves as another key benchmark.

### *Forecast of the trend LFP rate*

There are two additional issues in forecasting the trend LFP rate beyond 2007.

First, some birth cohorts had not reached one or more of our age groups by 2007. For example, no one born in 1995 was of legal age to work by 2007, implying that we cannot estimate a cohort effect for that birth year. Similarly, because our models are estimated separately for 25–54 year olds and 55–79 year olds, using data only through 2007, we have no estimated cohort effect for those born in, say, 1960 to determine their participation when they reach age 55 in 2015.

To overcome the lack of estimates for some birth cohorts, we forecast their cohort coefficients using a linear time trend over the last ten birth year coefficients. In other words, we project that cohort effects will slowly evolve in the future in the same way that they have over the previous decade. We do this separately for each sex, education level, and age group combination.

The idea is illustrated in figure 8, panels A, B, and C. The three panels plot the coefficients on the cohort dummies $\beta_{seb}$ for unmarried white women aged 25–54 without a child under five years old, by education level (high school dropout, high school graduate, and college graduate). The dashed lines at the end are the projections.

In addition to age and birth cohort dummies, our model also includes other time-varying (or age-varying) covariates (for example, family structures during prime working age or life expectancy at older ages). Therefore, the coefficients on the birth year dummies alone do not present a full cohort profile, but rather a profile conditional on specific values of the covariates. For example, the coefficients plotted in panels A, B, and C of figure 8 show the average differences across birth cohorts in (the log odds of) labor force participation for prime-working-age white women who are unmarried and have no young child, by educational attainment. Thus, the figure illustrates the evolution of cohort effects over time for this particular demographic group of women.[30]

Starting with those born in the late 1920s and continuing unabated for about four decades, newer birth cohorts were more likely to participate in the labor force during their prime working years, regardless of education level. This pattern reflects the dramatic increase in female labor force participation over the twentieth century (see figure 4, p. 105). However, for this group of women with at least a high school diploma (figure 8, panels B and C), that upward pattern reversed by the late 1960s. Cohorts with the same educational background born during the 1970s and early to mid-1980s were less likely to work than those born a decade or two earlier. Ultimately, this contributed to the flattening out of the female LFP rate by the mid-1990s and 2000s, offsetting the positive impact of higher female educational attainment on LFP throughout this period. By contrast, for women without a high school diploma, prime-working-age labor force participation has continued to slowly rise into the 1980s birth cohorts.[31]

Also of note, we do not attempt to estimate time-varying age effects (or other regression coefficients) in the forecasted trend. Instead, we simply apply the estimates obtained from the 1982–2007 sample to the simulated populations defined by age, education level, race, and sex (as well as by marital status and the presence of young children for certain age groups).[32]

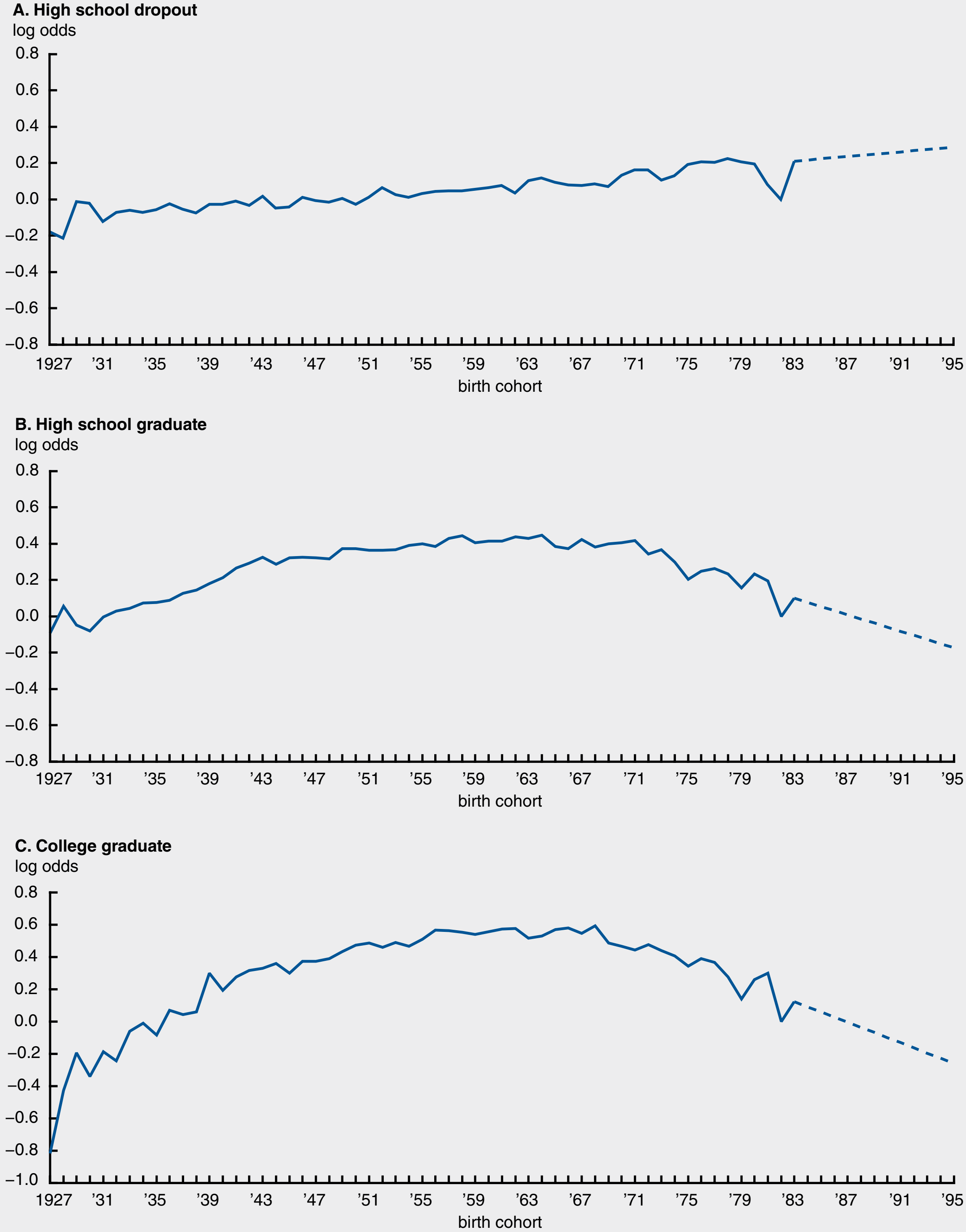
A second issue in making forecasts of LFP arises from the U.S. Census Bureau's forecast of the population. To forecast the trend LFP rate, we construct simulated populations for the rest of our forecast horizon using the civilian noninstitutional population projections by age, sex, and race prepared by the U.S. Census Bureau. But these projections are not broken down by education level. As a solution, we use a statistical model (see box 1) to predict educational attainment for age-sex-race groups, and then apply the LFP model to project the fraction of the people in these population groups that will be in the labor force.

The LFP model includes other age-group-specific demographic controls, such as marital status and the presence of young children for 25–54 year olds, which are, like education level, unavailable in the U.S. Census Bureau's population projections. Rather than build a model for these additional controls, we simply estimate the distribution of marital status and presence of young children within each age-sex-race group from the 2014 CPS and assume this distribution persists for the remainder of our forecast horizon. We similarly assume the unemployment gap, minimum wage, and youth-to-adult wage ratio return to their averages. We also assume life expectancy follows projections made by the Social Security Administration.[33]

These procedures are updates and extensions of Aaronson and Sullivan (2001). Aaronson et al. (2006) and Aaronson et al. (2014) follow a similar methodology. Compared with the methodology of these two papers, the main differences are as follows: 1) We estimate the

**FIGURE 8**

**Birth year coefficients for unmarried white females aged 25–54 without a young child, by education level**




Notes: The coefficients plotted in panels A, B, and C show the average differences across birth cohorts in (the log odds of) labor force participation for unmarried white females aged 25–54 without a young child (under five years old), by education level. See the text for further details. The solid lines represent cohorts that are fully seen in the age range 25–54, while the dashed lines represent projections.
Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*.

model at the individual level rather than at the age-sex group level; 2) we estimate the model conditioning on completed schooling, thus allowing demographics to affect LFP differently by education level; 3) we use the state unemployment gap to control for business cycle conditions at more geographically detailed labor markets; and 4) we allow the dynamic cyclical relationship between LFP and unemployment to vary by education level and demographics. These differences are detailed in the preceding discussion.

## Results

In this section, we discuss the results from our LFP models. We go over the baseline aggregate results first. Then we analyze the sources of changes in the trend LFP rate by examining the LFP of specific demographic groups; while doing so, we distinguish between shifts in the share of particular age groups and changes in certain groups' LFP behavior. Next, we study the gap between the actual LFP rate and our estimated trend LFP rate, first by educational attainment and then by age. Finally, we discuss the robustness of our results.

### *Baseline aggregate results*

As previously noted, figure 1 (p. 102) plots our measure of the LFP rate from CPS data (solid green line) against our baseline estimate of the long-run trend LFP rate (solid red line) between 1982 and 2020. Starting in the early 1980s, the trend rate of labor force participation rose uninterrupted through 2000. Since 2000, it has been falling—and at a steeper pace than that of the ascent. Consequently, as of the third quarter of 2014, our baseline model estimates the trend LFP rate to be 64.2 percent, almost 1 percentage point below its estimated level in the first quarter of 1982.

By construction, the trend LFP rate removes the effects of the business cycle by setting the unemployment gap to zero. As such, the gap between the actual and trend LFP rates is typically positive in periods such as the late 1990s, when the economy is growing rapidly, wage growth is strong, and more individuals are, therefore, willing to work than we might expect given the composition of the working-age population. By contrast, a negative LFP gap appears during recessions and weak recoveries. Indeed, we estimate that the negative LFP gap in 1982 was large (about –1.1 percentage points) and took much of the 1980s expansion to eliminate.

Similarly, during the most recent business cycle, the LFP gap was positive at the end of the last expansion in 2007. But the actual LFP rate fell more rapidly than the (declining) trend LFP rate, causing the LFP gap to turn negative during 2009, where it has remained since. As of the third quarter of 2014—over five years after the official end of the 2008–09 recession according to the National Bureau of Economic Research—the participation rate remains 1.2 percentage points below the long-run trend. This is much larger than our estimate of the LFP gap in 1988, just over five years after the end of the 1981–82 recession.

We project that the trend LFP rate will continue to fall by about 0.3 percentage points annually through at least 2020, at which point it will be 62.3 percent. The last time the actual LFP rate was that low was in the mid-1970s.

**BOX 1**

**Educational attainment models**

The U.S. Census Bureau's population data that we use to compute population weights in order to forecast the trend LFP rate are only available by sex, race, and age. Therefore, we follow the statistical model described in Aaronson and Sullivan (2001) to predict educational attainment.

Let $p_{it}^{j} = \text{Prob}\{y_{it} = j \mid j = 1, \ldots, 5\}$ be the probability that individual $i$ in year $t$ has education level $j$, where $j$ = 1 is high school dropout, $j$ = 2 is high school graduate, $j$ = 3 is some college, $j$ = 4 is college graduate, and $j$ = 5 is postcollege degree, and let $q_{it}^{j} = \text{Prob}\{y_{it} \geq j \mid y_{it} \geq j-1\}$, $j$ = 2, …, 5 be the probability that an individual reaches at least level $j$ given that he reached level $j$ – 1. We fit a statistical model to predict the $q_{it}^{j}$ and recover $p_{it}^{j}$ from $p_{it}^{j} = \prod_{k=2}^{j} q_{it}^{k}\left(1 - q_{it}^{j+1}\right)$. The $q_{it}^{j}$ is predicted based on a logistic regression model similar to the LFP models:

$$\log\left(\frac{q_{sbai}^{e}}{1 - q_{sbai}^{e}}\right) = \alpha_{sa}^{e} + \beta_{sb}^{e} + w_{(a+b)i}\lambda_{s}^{e} + x_{sbai}\gamma_{s,}^{e}$$

where the right-hand-side variables include indicators for age $\alpha_{sa}^{e}$, birth year $\beta_{sb}^{e}$, and race $x_{sbai}$, as well as business cycle controls $w_{(a+b)i}$.[1] The model is estimated separately by sex and race.[2]

[1]In all models, we also include an indicator variable for post-1992. This is because there was a redesign of the education question in the CPS in 1992, which causes a discrete change in some of the education categories. The parameters $\lambda_{s}^{e}$ and $\gamma_{s}^{e}$ are the regression coefficients on the $w$ and $x$ variables, respectively.

[2]The model for postcollege degree includes race as a control and generates estimates separately by sex only. Note that we impose different minimum age restrictions for each education-level-specific model to acknowledge that higher education levels begin at later ages. In particular, the high school graduate model includes only those aged 17 and older. Likewise, the some college, college graduate, and postcollege degree models include only those aged 19 and older, 21 and older, and 25 and older, respectively.

The dashed red line in figure 1 plots a prediction of the LFP rate that uses the contemporaneous state unemployment gaps (and their lags). This measure, which we label the "LFP rate prediction based on unemployment," reveals whether the actual LFP rate is unusually high or low given the present labor market situation. For example, during the late 1990s, the actual LFP rate was running above not only our estimate of the trend rate but also what we would have expected given the tight labor markets at the time. During the most recent recession and the ensuing expansion up through late 2013, the predicted LFP rate that accounts for contemporaneous economic conditions fell about the same as the actual LFP rate data (see figure 1). But since the fourth quarter of 2013, the actual LFP rate has fallen sharply, while our LFP rate prediction based on unemployment has not.[34] As of the third quarter of 2014, the actual LFP rate is 0.8 percentage points below where we would have expected given the unemployment rates that have prevailed over the past few years, suggesting there is significant slack in the labor market beyond that signaled by the unemployment rate.

***A decomposition of the trend LFP rate***

Next, we unpack the sources of changes in the trend LFP rate over the past 30 years into two parts: that due to changing demographics, in particular age and sex (which we call "demographic"), and that due to changing participation decisions *within* a given demographic group (which we call "behavioral"). The latter includes changes in some observed characteristics, such as education level, fertility rate, and life expectancy, as well as changes in unobservables captured by the cohort dummies.

In particular, let $p_t$ be the overall trend LFP rate at time $t$, $p_{dt}$ be the trend LFP rate for demographic group $d$ at time $t$, and $f_{dt}$ be the share of the population in group $d$ at time $t$. We can write the aggregate trend LFP rate as the weighted average of group-specific trend LFP rates,

$$p_t = \Sigma_d f_{dt} p_{dt},$$

and the change in the aggregate trend LFP rate as the sum[35]

$$\Delta p_t = \Sigma_d \left( p_{dt-1} - p_{t-1} \right) \Delta f_{dt} + \Sigma_d f_{dt} \Delta p_{dt}.$$

The first term on the right-hand side reflects the contribution from changing demographics ($\Delta f_{dt}$). An important recent example of $\Delta f_{dt}$ is the changing share of workers in their sixties. Since the standard life-cycle pattern suggests that those in their sixties work less than the aggregate working-age population (that is, $p_{dt-1} - p_{t-1} < 0$), the aggregate trend LFP rate declines (that is, $\Delta p_t < 0$) when the population share in their sixties increases and the trend rate rises when the population share in their sixties declines. The second term reflects the contribution from changing behavior for a given demographic group ($\Delta p_{dt}$). If those in their sixties are working longer today than in the past (that is, $\Delta p_{dt} > 0$ ), the aggregate trend LFP rate will rise (that is, $\Delta p_t > 0$). Table 1 reports the results of this decomposition of the aggregate trend LFP rate (based on the baseline estimates), split further by age for the demographic contribution and by gender and age for the behavioral contribution.

The top row in panel A of table 1 shows the annualized change in the aggregate trend LFP rate, reported over different subperiods (1982–97, 1997–2007, 2007–14, and 2014–20). The 1980s and 1990s were an era of rising LFP, and this is reflected in the increases of 0.11 percentage points per year in our trend LFP rate during 1982–97. Changing demographics (table 1, panel A, second row) explain a small part of this gain. Behavioral changes (table 1, panel A, third row), especially among prime-working-age women, play a more important role. In particular, rising LFP among women on account of behavioral factors contributed 0.17 percentage points per year to the change in the aggregate trend LFP rate over the 1982–97 period (table 1, panel C, sixth row). However, men's falling LFP throughout the 1980s and 1990s (table 1, panel C, second row) offset about half of these positive developments.

The tide began to turn around the turn of the century. After decades of increasing LFP, the trend LFP rate declined by 0.08 percentage points per year between 1997 and 2007 (table 1, panel A, first row). We attribute most of this decline to demographics (table 1, panel A, second row), as the oldest baby boomers hit their late fifties and began to exit the labor force (see also table 1, panel B, fourth row). While behavioral changes were virtually a neutral contributor (table 1, panel A, third row), that masks several continuing stories: increases in prime-working-age and older female participation (table 1, panel C, eighth and ninth rows) offsetting declines in prime-working-age male participation (table 1, panel C, fourth row) and in youth participation among both men and women (table 1, panel C, third and seventh rows). Falling youth LFP for both genders on account of behavioral factors contributed a total of –0.09 percentage points per year to the change in the aggregate trend LFP rate during 1997–2007.

Since around 2007, both the behavioral and demographic patterns have intensified, with the trend LFP rate falling by roughly 0.3 percentage points per year (table 1, panel A, first row). Demographic factors,

**TABLE 1**

**Decomposition of the trend labor force participation percentage change per year over subperiods of 1982–2020**

| | 1982–97 | 1997–2007 | 2007–14 | 2014–20 |
|---|---|---|---|---|
| **A. Decomposition of the trend percentage change per year** | | | | |
| Total change | 0.11 | −0.08 | −0.33 | −0.27 |
| Demographic | 0.03 | −0.07 | −0.29 | −0.22 |
| Behavioral | 0.09 | −0.01 | −0.05 | −0.04 |
| **B. Decomposition of demographic contribution to trend percentage change per year, by age** | | | | |
| Total demographic | 0.03 | −0.07 | −0.29 | −0.22 |
| Age 16–24 | −0.02 | 0.00 | 0.03 | 0.01 |
| Age 25–54 | 0.07 | −0.06 | −0.12 | −0.04 |
| Age 55 and older | −0.02 | −0.02 | −0.20 | −0.19 |
| **C. Decomposition of behavioral contribution to trend percentage change per year, by sex and age** | | | | |
| Total behavioral | 0.09 | −0.01 | −0.05 | −0.04 |
| Male | −0.08 | −0.06 | −0.07 | −0.03 |
| Age 16–24 | −0.03 | −0.05 | −0.07 | −0.04 |
| Age 25–54 | −0.04 | −0.05 | −0.05 | −0.03 |
| Age 55 and older | −0.01 | 0.04 | 0.05 | 0.04 |
| Female | 0.17 | 0.05 | 0.03 | −0.02 |
| Age 16–24 | 0.01 | −0.04 | −0.06 | −0.04 |
| Age 25–54 | 0.12 | 0.02 | 0.00 | −0.04 |
| Age 55 and older | 0.04 | 0.08 | 0.08 | 0.06 |

Notes: The estimated values shown are the annualized percentage changes in the trend rate of labor force participation based on data through 2007. The columns in each panel may not total because of rounding. See the text for details on demographic and behavioral contributions.
Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*.

especially the baby boomers reaching their fifties and sixties, explain about 90 percent of this decline; but behavioral factors are important as well (table 1, panel A, second and third rows). Continuing declines in prime-working-age male LFP and youth LFP (table 1, panel C, fourth and third and seventh rows) put downward pressure on the aggregate LFP rate, without the offsetting positive influences from prime-working-age female LFP (table 1, panel C, eighth row) that drove much of the gains in the late twentieth century. Older men and women are working more than in the past; but to date, the magnitude has been too small to offset other behavioral patterns (table 1, panel C, fifth and ninth rows).

Over the remainder of the decade, we expect many of these trends to continue. The trend LFP rate will continue to fall by nearly 0.3 percentage points per year. Demographics will still account for most (80 percent) of the decline, as baby boomers reach their seventies. In fact, the increase in the population share of those aged 70 and older alone will account for roughly half of the decline in the aggregate trend LFP rate (not shown in table 1). This occurs notwithstanding our expectation that the baby boomers will be working far longer than any past generation.

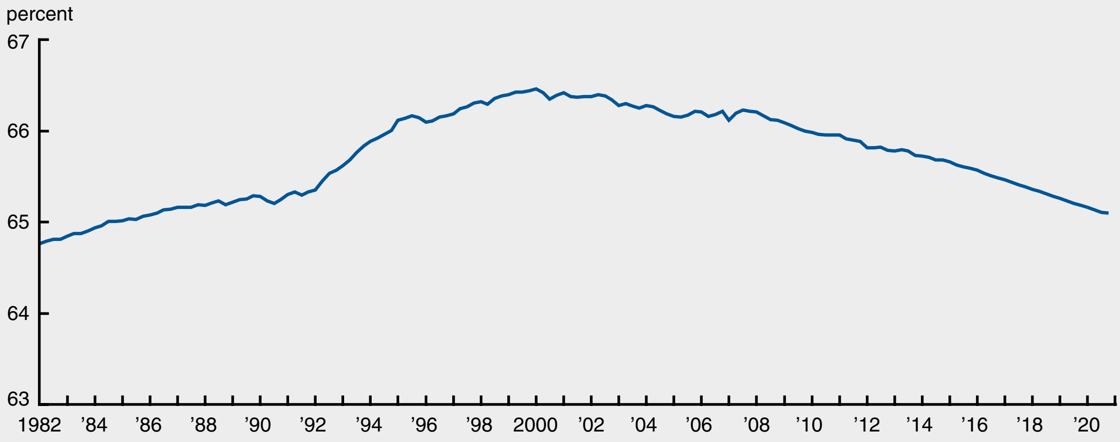
While evolving demographics (particularly those related to the large baby boom generation approaching or entering retirement) have been the focus in much of the recent discussion on the decline in LFP, we want to emphasize that long-running secular changes in work participation decisions within demographic groups have been an important part of the story as well. To show this more clearly, we plot in figure 9 a trend LFP rate that holds the age-sex groups' population shares fixed at their 2007 levels but allows the group-specific trend LFP rates to vary over time as predicted by our model. This demographically adjusted hypothetical trend LFP rate is still moving down between 2007 and 2013, highlighting that there are factors besides an aging population at play. Next, we turn to some of these specific patterns in more detail.

### *LFP gap by education level*

Figure 10 plots the actual and trend LFP rates for those aged 25 and older by education level.[36] Between the late 1990s and 2007, the actual and trend LFP rates moved steadily down for those with at least a high school diploma. The LFP of high school dropouts was the exception (figure 10, panel A).[37] However, since 2007, the actual LFP rate of high school dropouts has stopped

FIGURE 9

Demographically adjusted trend labor force participation (LFP) rate, 1982–2020




Note: The figure plots a trend LFP rate that holds the population shares fixed at their 2007 levels by single age and sex while allowing the group-specific trend LFP rates to vary over the period 1982:Q1–2020:Q4.
Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*.

increasing and has even fallen a little, opening up a large gap between itself and our estimated trend LFP rate. As of the third quarter of 2014, the actual high school dropout LFP rate is about 2.5 percentage points below where we would expect given other demographic characteristics and a neutral labor market. Since 2007, the actual LFP rates have fallen for groups with higher educational attainment as well. However, these declines were better anticipated by long-running demographic patterns within these groups. For example, as of the third quarter of 2014, the LFP gap is about –1.1 percentage points for high school graduates (figure 10, panel B) and essentially zero (+0.2 percentage points) for college graduates (figure 10, panel D). Indeed, for the latter group, a significant LFP gap never materialized throughout the recent recession and slow recovery.

A similar pattern emerges when we measure the gap between the actual LFP rate and our LFP rate prediction based on unemployment (whose aggregate measure is featured in figure 1 on p. 102 but which is not shown in figure 10). This predicted measure takes into account the high unemployment in recent years. As of the third quarter of 2014, the LFP gap between the actual rate and this predicted rate based on unemployment is –1.4 percentage points for high school dropouts and –0.7 percentage points for high school graduates but +0.3 percentage points for college graduates. That is, given the recent labor market conditions and demographic characteristics, a surprising share of workers without a college degree have dropped out of the labor force since 2007.

Why an LFP gap has opened up for workers without a college degree is of significant policy interest. One interpretation is that it reflects an extra measure of labor market slack not reflected in unemployment rates. However, it is also possible that some of the gap may reflect new but potentially long-running phenomena not captured by our model. For example, middle-income-paying jobs, often in manufacturing, that in the past could have been filled by less educated workers are disappearing (Acemoglu and Autor, 2011, and the references therein). Workers who traditionally have filled those occupations are being forced to adapt by taking on jobs that have traditionally been filled by low-skill workers, such as teens. That, in turn, has put significant wage pressures on the low-skill labor market, potentially pushing many to leave the labor force altogether. The 2000s housing boom may have temporarily stopped the slide of real wage rates of low-education workers (Charles, Hurst, and Notowidigdo, 2014a, 2014b), and thus temporarily held up the actual LFP rate, as well as our estimated trend rate, for low-education workers; but eventually, the housing collapse led to both wage and LFP rate declines.

Once a worker experiences a long spell of unemployment, it can be difficult to overcome. Employers may use length of unemployment as a signal of quality, and shun those who are unemployed beyond short-term spells (Blanchard and Diamond, 1994). A recent experiment reported in Kroft, Lange, and Notowidigdo (2013) indicates that callback rates are lower for those with

**FIGURE 10**

**Labor force participation rates for those aged 25 and older, by education level, 1982–2020**

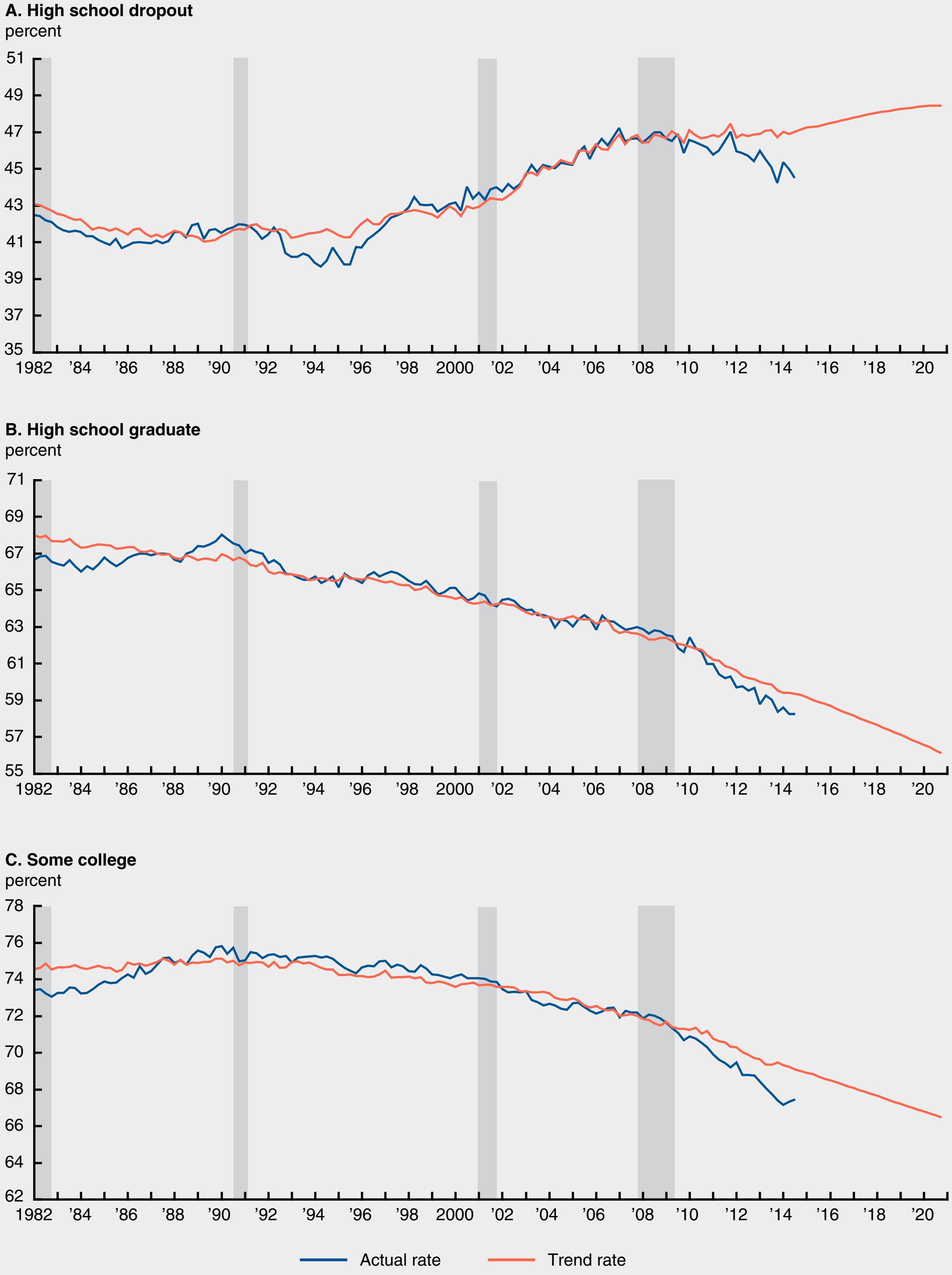
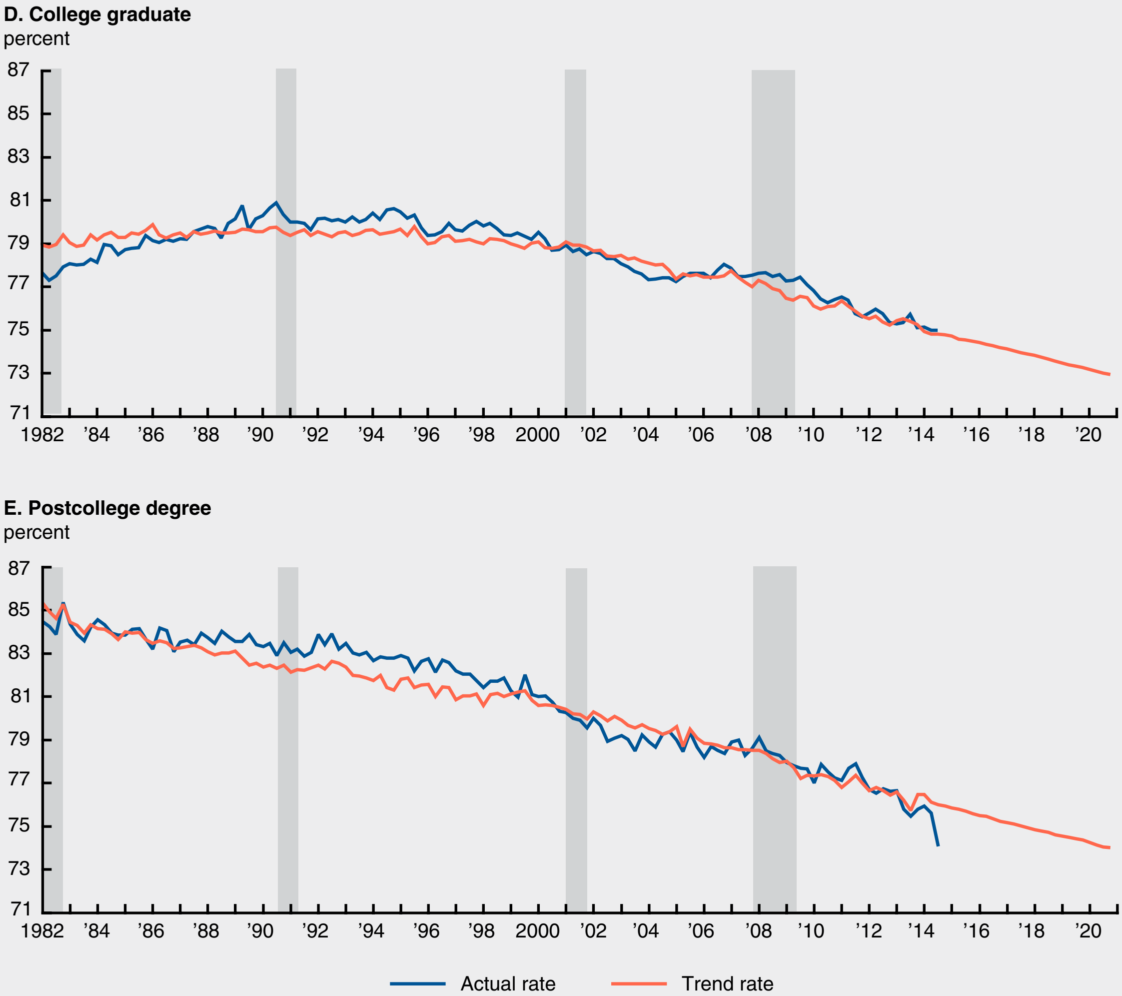


**FIGURE 10 (CONTINUED)**

**Labor force participation rates for those aged 25 and older, by education level, 1982–2020**




Notes: Each panel plots quarterly data for those aged 25 and older over the period 1982:Q1–2020:Q4. The shaded bars indicate recessions as defined by the National Bureau of Economic Research.
Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*.

longer ongoing unemployment spells, conditional on other aspects of a resume that employers value.

It is also possible that the decline in LFP among low-education workers is related to social safety net programs—in particular, the Social Security Disability Insurance program. DI rolls have been increasing throughout the most recent business cycle, continuing a pattern that has been more or less uninterrupted since the 1990s (Autor, 2011; and Burkhauser and Daly, 2011). DI tends to be countercyclical partly because eligibility standards ease amid deteriorating labor market conditions (Mueller, Rothstein, and von Wachter, 2013). That is, people with moderate disabilities are more likely to qualify for the program when there are fewer suitable jobs available.

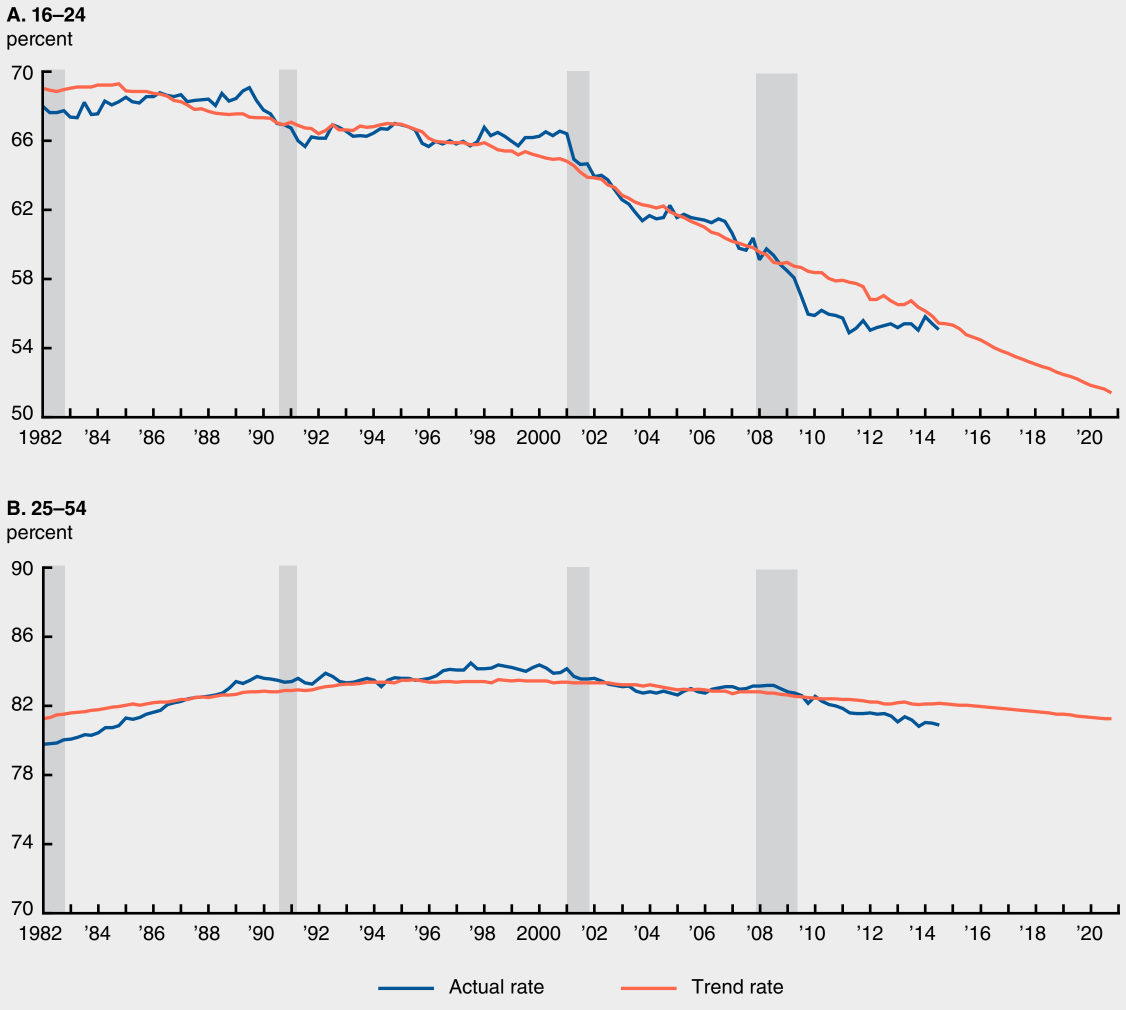
Finally, the expected upward trend in LFP of those without high school diplomas may have been driven by the welfare reforms of the 1990s, when policy induced more low-education women to work. That policy intervention may have been interpreted by the model as a trend that would continue rather than as a one-time change to the level of LFP.

### *LFP gap by age*

Figure 11 plots the actual and trend LFP rates by age. A sizable gap between the actual LFP rate and our estimated trend LFP rate opened up among 16–24

FIGURE 11

Labor force participation rates, by age, 1982–2020




year olds during the Great Recession and the early part of the subsequent recovery, but that gap has largely closed. Today, the negative LFP gap is concentrated among prime-working-age workers (figure 11, panel B).

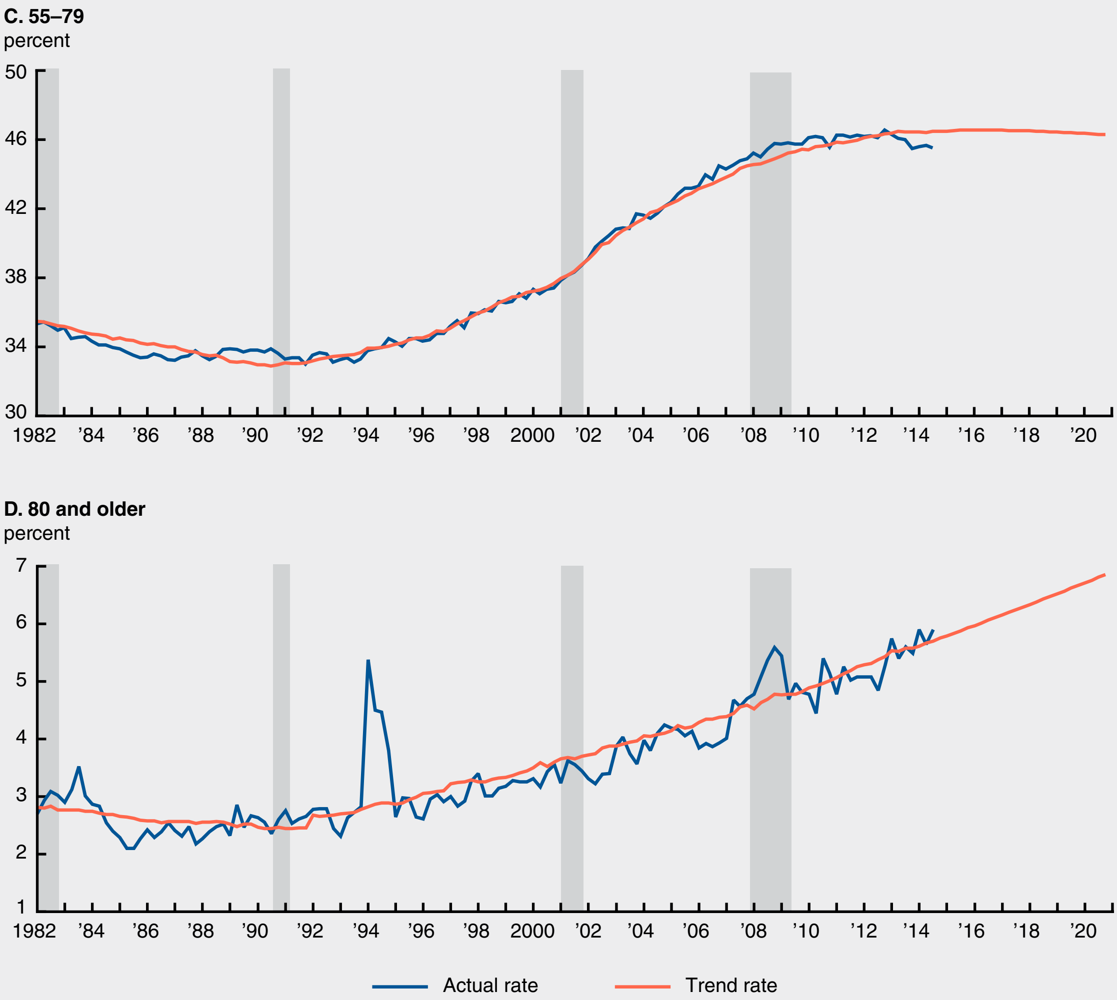
### *Robustness of results*

We experimented in a number of ways to gauge the robustness of our estimated trend LFP rate to reasonable alternative specification and measurement choices. Table 2 summarizes some of these exercises.

Estimating our model with data through the third quarter of 2014 rather than 2007 has little impact (a difference of about 0.2 percentage points) on our current estimates of the trend LFP rate and the LFP gap between the actual and trend rates (compare the first and second rows of table 2). We also estimate our model through the final quarter of 2002, 2004, and 2006 and the results remain similar—the estimates for the LFP gap in the third quarter of 2014 are –1.1, –0.4, and –0.9 percentage points, respectively. It is worth emphasizing that our estimate of the trend LFP rate has remained quite stable since 2002 (see dashed lines in figure 12, panel A). This stands in stark contrast to the BLS's estimate of the trend LFP rate (Toossi, 2004, 2005, 2007), which panel B of figure 12 shows has changed considerably over time (dotted lines). The robustness of our results (as demonstrated by the dashed trend LFP rate lines all heading similarly lower in figure 12, panel A) reflects our methodology, which extrapolates labor force participation decisions from specific birth cohorts.

**FIGURE 11 (CONTINUED)**

**Labor force participation rates, by age, 1982–2020**




Notes: Each panel plots quarterly data over the period 1982:Q1–2020:Q4. The shaded bars indicate recessions as defined by the National Bureau of Economic Research.
Source: Authors’ calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*.

Using alternative measures of labor market tightness such as the national unemployment gap (table 2, third row), using a different lag structure (unreported), or adding the median duration of national unemployment (table 2, fourth row) had a small impact, altering our current estimate of the trend LFP rate by at most 0.3 percentage points.

That said, our estimate of the trend LFP rate is relatively sensitive to one critical modeling choice. The baseline model stratifies the estimation sample into four age groups (16–24, 25–54, 55–79, and 80 and older). However, when we estimate a model that forces cohort coefficients to be the same for ages 16–79 (the pooled model), we find that the trend LFP rate is almost identical to the one estimated from the baseline model through 2007, but diverges appreciably from then on. As of third quarter of 2014, the aggregate trend LFP rate estimated from the pooled model is 1.0 percentage point lower than that estimated from the baseline model. Thus, the LFP gap between the actual rate and the trend rate estimated from the pooled model is relatively smaller, at about –0.2 percentage points (table 2, fifth row).[38]

The differences between the trend and gap estimates in the first row and the fifth row of table 2 can be partly explained by how each model estimates and extrapolates the coefficients for the cohorts who were born too late to appear in the age-group estimation samples. The baseline model estimates the cohort effects separately by three age groups and extrapolates the future cohorts based on the evolution of the recent cohorts when they were at *the same age*. In contrast, the pooled model estimates a single set of cohort effects for all ages 16–79. This reduces the number of cohort effects that needs to be forecasted, but it imposes strong restrictions on the data. One problem with the pooled model approach is that data for the 16–24 year olds might not be very informative about LFP later in people's careers. This is a particular concern for high school dropouts. Many of those without a high school diploma at young ages will go on to get a diploma and thus won't be a good benchmark for older high school dropouts. However, as it turns out, when we pool the 16–24 and 25–54 year old samples together, the estimated aggregate trend LFP rate (table 2, sixth row) is not very different from that estimated from the baseline model.

Instead, the divergence between the estimates of the trend LFP rates from the baseline and pooled models arises from how we handle the older population. To derive the results shown in the final row of table 2, we combine the age 25–54 sample with the age 55–79 sample and find that the estimate of the trend LFP rate is similar to that from the pooled model (table 2, fifth row). We interpret this to mean that if the cohort effects impact labor force participation differently over the life cycle, then restricting them to be constant across all ages, as in the pooled model, could lead to misleading inferences.

To illustrate this point, we generated panel A of figure 13, which compares the coefficients on the birth year dummies from the baseline and the pooled models for unmarried white male high school dropouts without a young child (under five years old). The pooled model (solid green line) suggests that birth cohort effects have been fairly stable for much of the twentieth century. In the baseline model, however, the cohort effects exhibit different patterns at the three stages of the life cycle. Similar to cohort effects of the pooled model, cohort effects in the baseline model are relatively stable during youth and prime working age (purple and orange lines, respectively). By contrast, the birth cohort coefficients are rising quickly over time for the 55–79 age group (solid red line), indicating a notable difference in the likelihood of working past age 54 for those born at the beginning versus the middle of the twentieth century—a trend that we expect to continue for cohorts born later in the century (dashed red line).

Since the models also include other time-varying (or age-varying) covariates, the cohort effects alone do not give a full picture of the differences in the likelihood of labor force participation. So, in panel B of figure 13, we compare the predicted LFP rates of individuals of different ages in the third quarter of 2014 from the two models. As expected, while the two models yield similar predictions for prime-working-age workers (for example, those born in 1970), the pooled model predicts much lower labor force participation for individuals aged 55 and older (for example, the 1950 birth cohort) than the baseline model.

In principle, since the pooled model can mask important changes in cohort effects, we prefer the more flexible baseline model. Moreover, formal statistical tests also favor our baseline model over the pooled model, which is more restricted. In particular, we can reject the null that the cohort coefficients are the same across ages (that is, the restriction imposed by the pooled model) at the 1 percent level for all ten sex-education-level variations of the baseline model that we estimate.[39]

To summarize, while there is some uncertainty about the exact size of the current LFP gap, we view the robustness exercises as confirming that a significant part (but not all) of the decline in LFP rate since 2000—and since 2007—can be explained by changing demographic and behavioral factors. Relative to the results from two recent and related *Chicago Fed Letter* articles (Aaronson, Davis, and Hu, 2012; and Aaronson and Brave, 2013), the magnitude of the

**TABLE 2**

**Trend rate of labor force participation (LFP) and LFP gap in 2014:Q3**

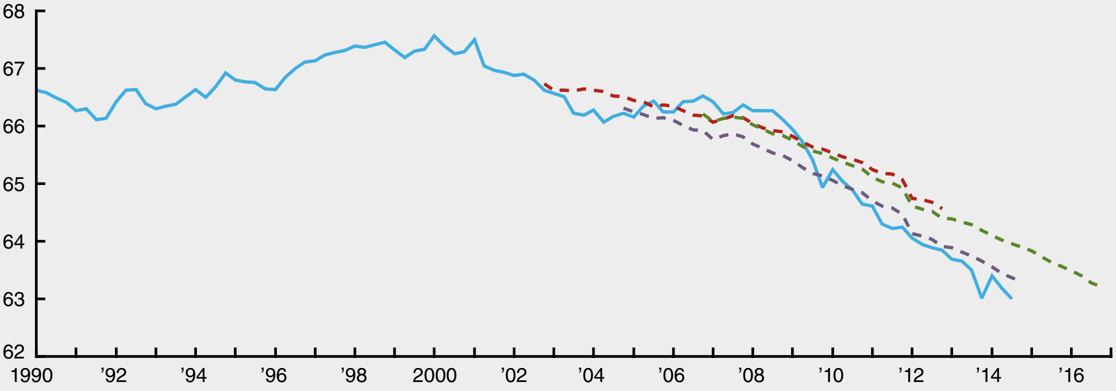
| Model change | Trend LFP rate | LFP gap |
|---|---|---|
| | *(percent)* | *(percentage points)* |
| Baseline | 64.2 | –1.2 |
| Estimate model through 2014:Q3 | 64.0 | –1.0 |
| National unemployment gap | 64.0 | –1.0 |
| Median duration of national unemployment | 64.5 | –1.5 |
| Pooled model | 63.2 | –0.2 |
| Age groups 16–54, 55–79, and 80 and older | 64.0 | –1.0 |
| Age groups 16–24, 25–79, and 80 and older | 63.0 | 0.1 |

Notes: The LFP gap is the difference between the actual and trend rates of LFP. The pooled model forces cohort coefficients to be the same for ages 16–79 (instead of differentiated for ages 16–24, 25–54, and 55–79). See the text for further details on the baseline model and variations of this model.
Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*.

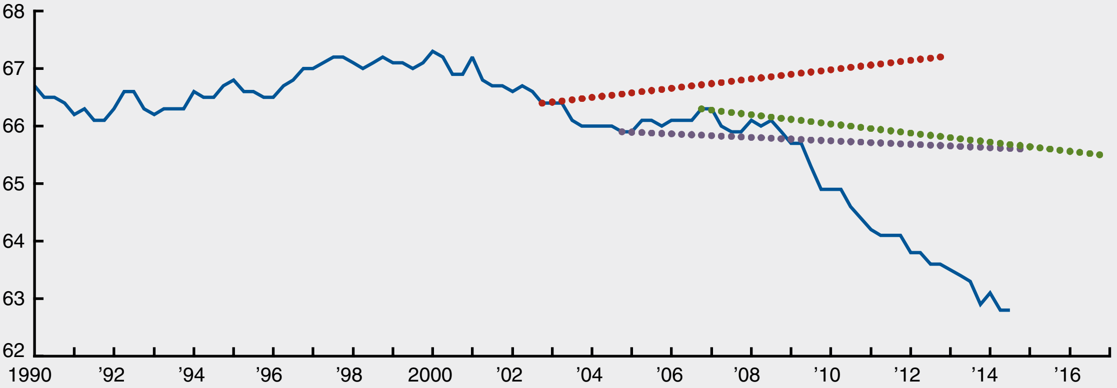
**FIGURE 12**

**Chicago Fed versus U.S. Bureau of Labor Statistics (BLS) labor force participation (LFP) rates and projections, 1990–2016**

**A. Chicago Fed model projections**




**B. BLS projections**

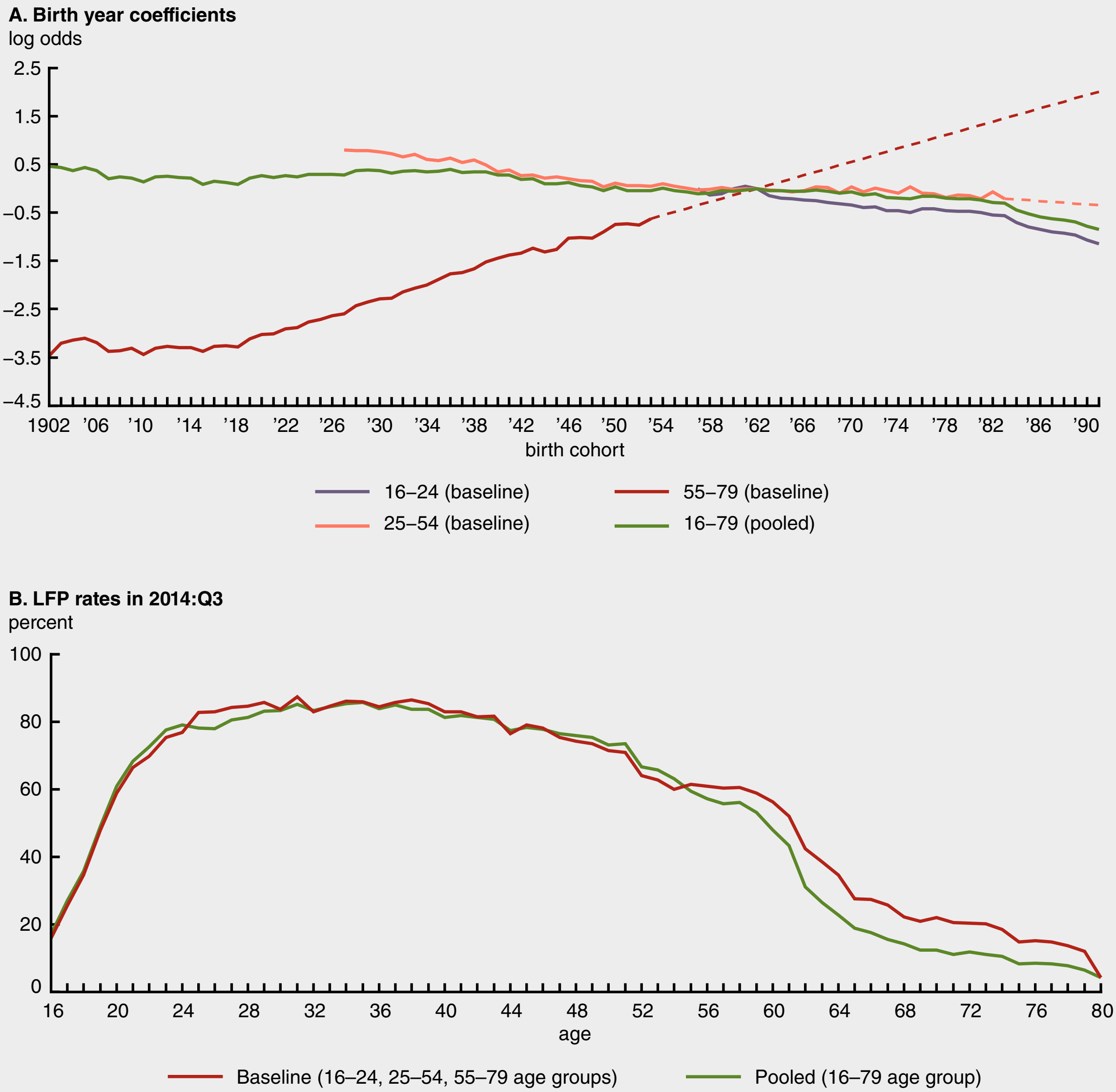



Notes: The figure plots quarterly data for those aged 16 and older over the period 1990:Q1–2016:Q4. The blue lines in both panels show actual LFP data; the light blue line in panel A is from the Chicago Fed's CPS calculations, while the dark blue line in panel B is from the BLS's CPS calculations. The difference between the actual LFP rates from the Chicago Fed and from the BLS is discussed in note 16. The dashed lines in panel A show the forecasted trend LFP rates from the Chicago Fed baseline model (described in the text) using data through different end dates. The dotted lines in panel B represent BLS forecasts starting at various dates.

Sources: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey* (CPS); and Toossi (2004, 2005, 2007).

**FIGURE 13**

**Labor force participation (LFP) of unmarried white male high school dropouts without a young child: Baseline model versus pooled model**

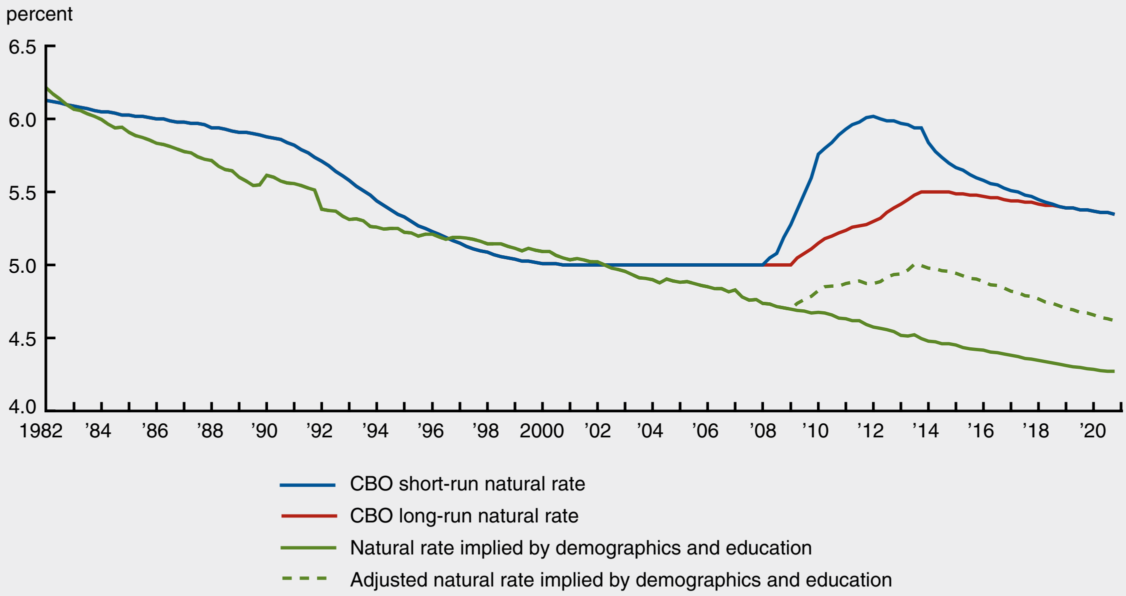



Notes: The pooled model forces cohort coefficients to be the same for ages 16–79 (instead of differentiated for ages 16–24, 25–54, and 55–79, as in the baseline model); see the text for further details. In panel A, the solid lines represent cohorts that are fully seen in those age ranges, while the dashed lines represent forecasts. A young child is a child aged under five years old.
Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*.

baseline LFP gap in 2011 (the latest comparable period) is somewhat smaller in this article. This change can be explained by data and modeling improvements. First, we now use real-time BLS population estimates that were released with the CPS data. In the previous two studies, we used the resident population estimates that were released by the U.S. Census Bureau in 2008. Second, we now reference the civilian noninstitutional population rather than the total population, to be consistent with published BLS figures. Finally, we have made a number of modeling improvements, including using higher-frequency data (quarterly versus annual), using fewer age groups, and allowing for long lags of the state unemployment rate. In total, these data and modeling changes cut our estimate of the LFP gap in 2011 by a quarter of a percentage point—from just over 1 percentage point to about 0.8 percentage points. The gap has widened slightly since 2011.

FIGURE 14

Natural rates of unemployment, 1982–2020




Notes: The figure plots quarterly data for those aged 16 and older over the period 1982:Q1–2020:Q4. The natural rate implied by demographics and education is the natural rate of unemployment that holds the specific natural rates of unemployment for age-sex-education-level groups fixed at their values in the second half of 2005 while allowing group-specific shares of trend labor force participation to vary over time. For further details on the groups, see the text. The dashed green line is this natural rate implied by demographics and education adjusted with the same post-2007 adjustment as the CBO estimate for its long-run natural rate; see the text, particularly note 42, for further details.
Sources: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*; and Congressional Budget Office (CBO) from Haver Analytics.

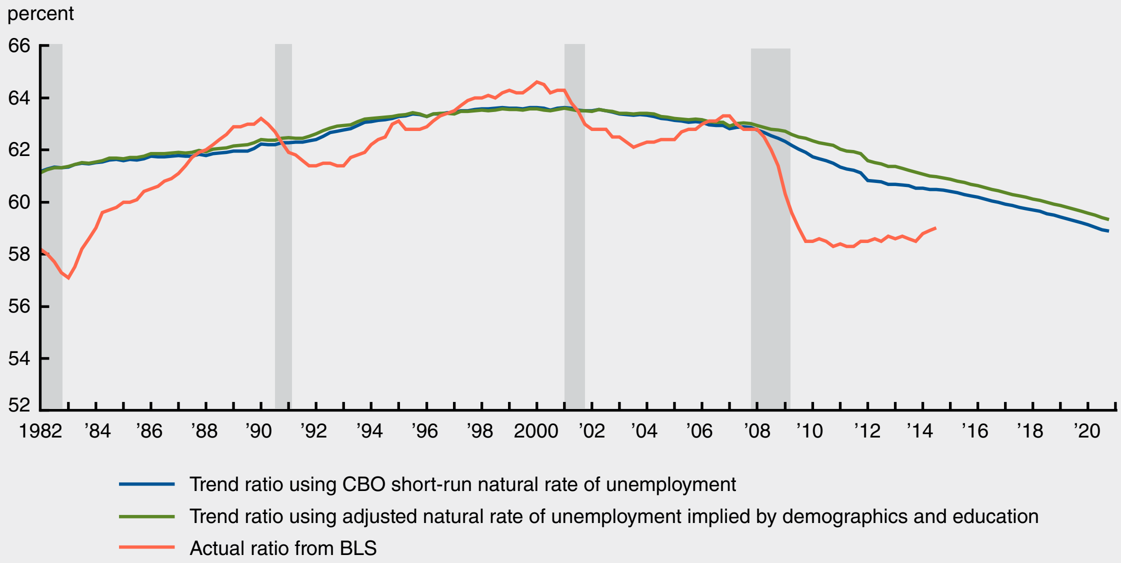
## Impact on the natural rate of unemployment

As we have documented, declining trend labor force participation is a widespread phenomenon, but the magnitude of the decline differs across demographic groups. A consequence of this heterogeneity in the decline of trend LFP is that the composition of the aggregate labor force has changed over time, which in turn can impact the natural rate of unemployment, or trend unemployment rate. For example, we estimate that the trend LFP rate has fallen especially rapidly for teens—a group that happens to have particularly high rates of unemployment today. As teens become a smaller share of the labor force, the natural rate of unemployment will decline. In addition, educational attainment has been increasing over time—a development that increases the share of workers with lower-than-average unemployment.

To broadly assess the likely magnitude of this compositional effect, we calculate (from CPS data) a trend unemployment rate implied by demographics and education—specifically, one that holds the specific trend unemployment rates for the age-sex-education-level groups fixed at their respective levels in the second half of 2005 (a time when the actual aggregate unemployment rate was equal to the Congressional Budget Office's estimate of the natural rate of unemployment) but allows these groups' shares of the trend LFP to vary over the entire period 1982–2020. As figure 14 shows, this hypothetical natural rate of unemployment rate (solid green line) declines by 0.3 percentage points over the period 2007–14 and by 0.6 percentage points over the period 2000–14, or about 0.05 percentage point per year over the past 15 years. In other words, the aggregate natural rate of unemployment is 0.3 percentage points lower in the third quarter of 2014 than it would have been if the composition of the trend LFP had remained the same as in 2007 and 0.6 percentage points lower than it would have been if this composition had remained the same as in 2000.[40] The decline since 1982 in the natural rate of unemployment implied by demographics and education is also quite similar to that of the CBO's natural rate of unemployment series (in figure 14, the CBO's short-run natural rate is the blue line, and its long-run natural rate is the red line[41]), though the timing is somewhat different. Both of the CBO's natural rate of unemployment series declined from 6.1 percent at the beginning of the series in 1982 and flattened to

FIGURE 15

**Employment-to-population ratios, 1982–2020**




Notes: The figure plots quarterly data for those aged 16 and older over the period 1982:Q1–2020:Q4. See the notes of figure 14 and the text for further details on the natural rates of unemployment. The shaded bars indicate recessions as defined by the National Bureau of Economic Research.
Sources: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey* (CPS); and Congressional Budget Office (CBO) from Haver Analytics.

5.0 percent from 2000 through 2007, whereas the trend unemployment rate implied by demographics and education declined steadily from about 6.2 percent in the early 1980s to 4.8 percent in 2007. According to the CBO's estimates, the short-run natural rate rose sharply during the most recent recession that began in late 2007, peaking at 6.0 percent in 2012, while the CBO's long-run natural rate increased more steadily, hitting 5.5 percent in 2013 before declining slightly to meet the short-run rate by the 2020 estimate of 5.4 percent (both CBO series have yet to return to pre-recession levels). When we apply the same post-2007 increase in the CBO long-run estimate of the natural rate to our hypothetical trend unemployment rate, this adjusted natural rate (dashed green line) currently stands at 5.0 percent.[42] In the next section, we use this adjusted natural rate implied by demographics and education, as well as the CBO's short-run natural rate, to calculate trend payroll employment growth.

Figure 15 shows the implication of our trend LFP and natural rate of unemployment results for the trend employment-to-population ratio.[43] As the figure shows, the trend employment-to-population ratio has been falling for over a decade because of the drop in the LFP rate. The value of the trend employment-to-population ratio using the CBO's short-run natural rate of unemployment (blue line) is 60.5 percent in the third quarter of 2014—about 1.5 percentage points greater than the actual BLS data (orange line). Relative to the trend employment-to-population ratio using the adjusted hypothetical trend unemployment rate described in the previous paragraph (green line), the actual ratio is 2 percentage points lower in the third quarter of 2014.

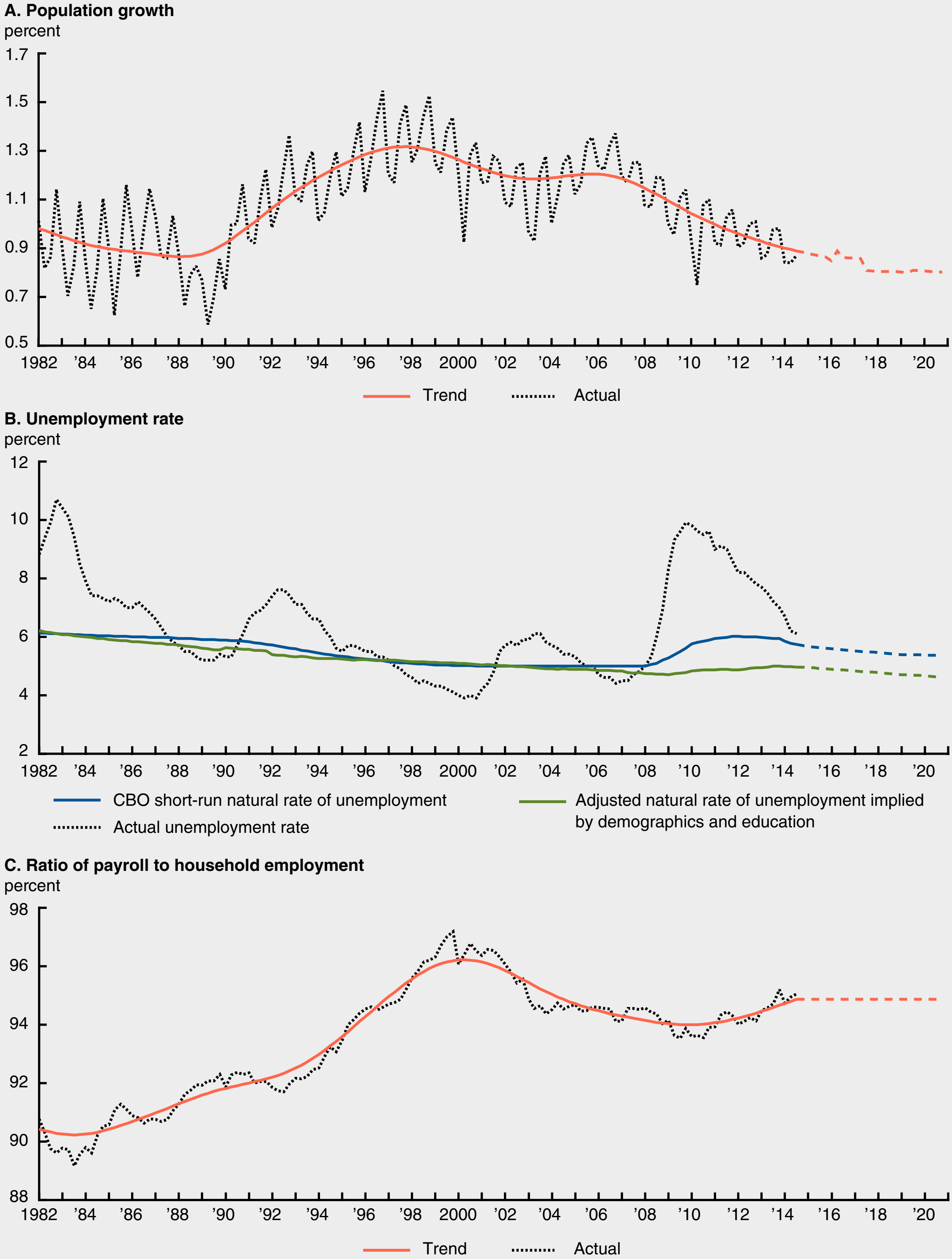
## Impact on trend payroll employment growth

In order to calculate trend payroll employment growth from 1982 through 2020, we use four estimated components: our baseline estimate of the trend labor force participation rate; the trend civilian noninstitutional population aged 16 and older; one minus the CBO's short-run natural rate of unemployment (or the adjusted natural rate of unemployment implied by demographics and education); and the trend ratio of payroll to household survey employment.[44] Trend payroll employment *growth* is the monthly average change implied by the product of these four constructed measures. Aaronson and Brave (2013) provide more supporting details.

Figure 16 plots the additional series needed for this calculation. Trend population growth[45] climbed steadily through the 1990s, peaking in the late 1990s at about 1.3 percent (figure 16, panel A). Trend population

**FIGURE 16**

**Underlying data series in trend payroll employment growth calculation**

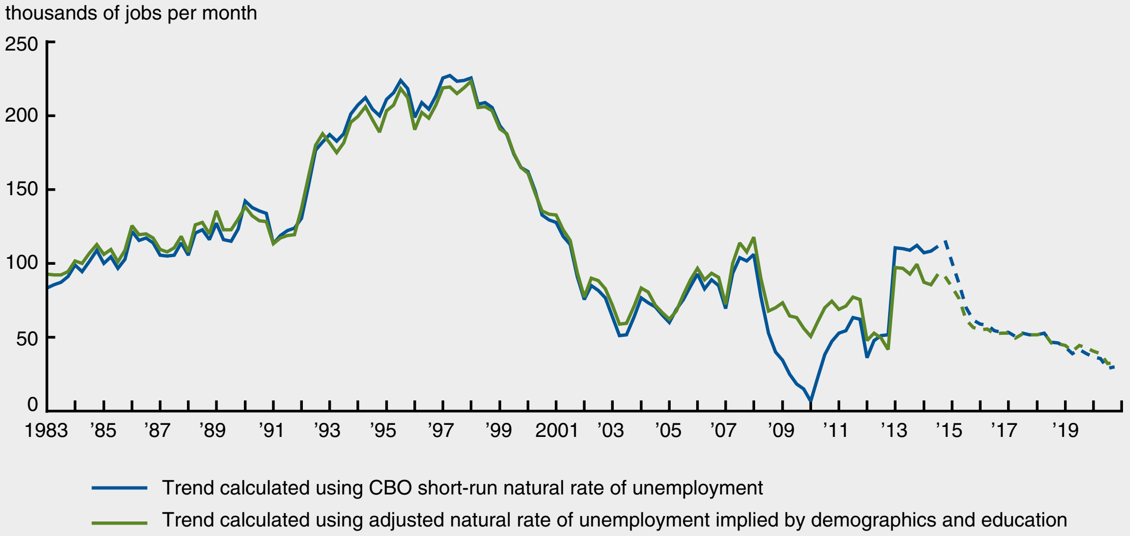



Notes: The figure plots quarterly data for those aged 16 and older over the period 1982:Q1–2020:Q4. The dashed lines in all three panels represent projections of the trends. Both natural rates of unemployment in panel B also appear in figure 14; see the notes of figure 14 and the text for further details.

Sources: Authors' calculations based on data from the U.S. Census Bureau; U.S. Bureau of Labor Statistics, *Current Population Survey* and *Current Employment Statistics*, from Haver Analytics; and Congressional Budget Office (CBO) from Haver Analytics.

FIGURE 17

Trend payroll employment growth, 1983–2020




Notes: The figure depicts the average monthly change in the trend in payroll employment on a quarterly basis. The data are smoothed using a four-quarter moving average. Solid lines denote estimates based on historical data, while dashed lines signify estimates based on projected data. See the notes of figure 14 and the text for further details on the natural rates of unemployment.
Sources: Authors' calculations based on data from the U.S. Census Bureau; U.S. Bureau of Labor Statistics, *Current Population Survey* and *Current Employment Statistics*, from Haver Analytics; and Congressional Budget Office (CBO) from Haver Analytics.

growth then decelerated to 0.9 percent in 2012, and the U.S. Census Bureau expects it to fall through 2016 and then stabilize at about 0.8 percent per year for the remainder of the decade. Panel B of figure 16 shows the CBO short-run natural rate of unemployment and the adjusted natural rate of unemployment implied by demographics and education discussed in the previous section, as well as the actual unemployment rate from the BLS. Finally, to derive an estimate of the trend in the more commonly referenced BLS payroll survey of employment requires an additional multiplication by the trend ratio of payroll to household survey employment.[46] The trend ratio of payroll to household employment recently stabilized at about 94.8 percent after a long ascent during the 1980s and 1990s and subsequent decline since 2000 (figure 16, panel C).[47] We expect it to stay at its current level of about 94.9 percent through 2020.

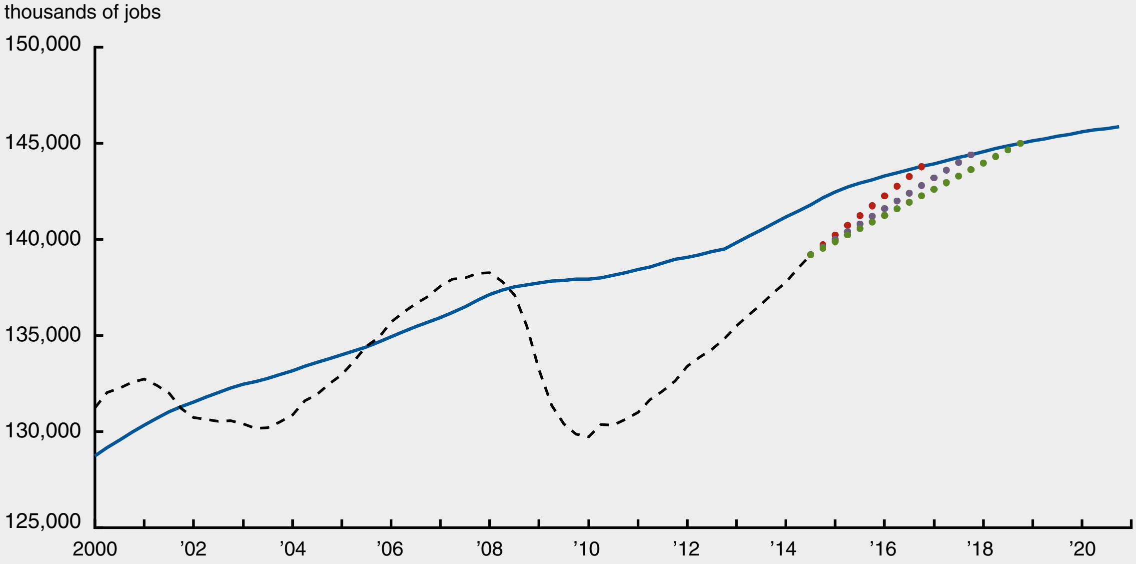
Figure 17 plots our estimate of trend payroll employment growth from 1983 through 2020[48] using both the CBO short-run natural rate of unemployment and the adjusted trend unemployment rate implied by demographics and education discussed before. Trend payroll employment grew by roughly 130,000 jobs per month during the mid-1980s through the early 1990s and by roughly 200,000 jobs per month during the middle to late 1990s. In the early 2000s, trend employment growth fell to under 100,000 jobs per month, where it has roughly remained.

The historically high rates of trend job growth in the 1980s and 1990s were driven by a confluence of all four factors described in this section—an increase in the trend labor force participation rate, higher trend population growth, a decline in the natural rate of unemployment, and an increase in the trend ratio of payroll to household survey employment. As discussed before, the trend LFP rate, along with the trend ratio of payroll to household survey employment, reversed course around the turn of the century, causing trend payroll employment growth to fall.

During the Great Recession, trend payroll employment growth fell substantially, driven by a sharp rise in the natural rate of unemployment (especially in the CBO's short-run natural rate). With trends in the unemployment rate turning down during the recovery, trend payroll employment growth has subsequently picked up, averaging roughly 60,000–70,000 jobs per month since the expansion started in June 2009. We project trend employment growth to continue at the 60,000–70,000 jobs per month pace through the end of 2015 and then drop to under 50,000 per month, on average, in 2016–20. The projected slowdown is based on the continuing decline in trend labor force participation, along with a lower level of projected population growth from now on.

FIGURE 18

Payroll employment gap, 2000–20




Trend
Actual
Projection based on average monthly job gains of approximately 170,000 (closing gap by end of 2016)
Projection based on average monthly job gains of approximately 130,000 (closing gap by end of 2017)
Projection based on average monthly job gains of approximately 115,000 (closing gap by end of 2018)

Notes: The figure plots quarterly data for those aged 16 and older over the period 2000:Q1–2020:Q4. The colored dotted lines are potential future paths for when the gap between the actual and trend levels of total payroll employment would be closed.
Sources: Authors' calculations based on data from the U.S. Census Bureau; U.S. Bureau of Labor Statistics, *Current Population Survey* and *Current Employment Statistics*, from Haver Analytics; and Congressional Budget Office (CBO) from Haver Analytics.

It is worth noting that the calculations here are for *trend* payroll employment growth. While the trend is expected to slow down substantially over the rest of the decade, a large employment gap (that is, the difference between the actual and trend level of total payroll employment) that opened up during 2008–09 needs to be closed by above-trend employment growth. To illustrate some potential future paths, we show in figure 18 that if payroll employment grows by roughly 130,000 jobs per month, it will take three years for the gap to completely disappear (in 2017). Stronger employment growth of roughly 170,000 jobs per month closes the gap one year earlier (in 2016), while weaker employment growth of roughly 115,000 jobs per month closes the gap one year later (in 2018).

## Conclusion

This article extends previous methodology for estimating the long-run trend in LFP as well as its dependence on business cycle conditions. We find that our methodology—which 1) takes into account the changing distribution of educational attainment and other characteristics of the population, 2) uses state variation in unemployment gaps to identify the sensitivity to labor market conditions, 3) accounts for the life-cycle pattern of LFP, and 4) allows for flexible variation by birth cohort in how the life-cycle pattern develops—is quite robust in its implication that the trend LFP rate is moving down by about 0.3 percentage points per year. Our baseline results use data through 2007 to estimate the trend in LFP. However, we get very similar predictions of the decline in the trend if we estimate the model using data through the third quarter of 2014 or limit the data to a date as early as 2002.

Of course, there are many questions that our statistical models are not designed to answer. We do not account for the detailed functioning of the many public policy programs that impact work decisions. We also do not model the underlying supply and demand for labor in a manner that would provide insight into how LFP and wages are jointly determined. As a result, it is possible that policy changes such as those pertaining to disability insurance or education programs or endogenous changes to wage growth could alter the path of the LFP rate in the years ahead. Future research that builds such structural representations of the supply and demand for labor could thus be very valuable. That said, the trends our model identifies have been stable for nearly 15 years, so their predictions may provide a reasonable benchmark for future research. They imply that once employment has returned to its long-run trend, it will grow much more slowly than in the past, with typical employment gains of under 50,000 per month. Estimates of potential output growth and the natural rate of unemployment should also reflect lower projections for LFP and changes in the composition of the work force.

There is a good deal of interest among policymakers in the gap between the current level of the LFP rate and its long-term trend because it has implications for the stance of monetary policy. With regard to how large the LFP gap is, our results suggest a somewhat wider range of possible answers. Clearly, a large portion of the decline in the trend LFP rate since 2007 reflects demographics and other long-running factors. However, plausible models imply gaps of between 0.2 and 1.2 percentage points, depending on various details, especially on how we treat cohort effects for different age groups. Our preferred model, which allows the cohort patterns to be different for older and younger people, estimates the current gap relative to expectations in 2007 at about 1.2 percentage points.

We estimate that much of today's gap between the actual LFP rate and its trend is accounted for by low-education workers, possibly reflecting the especially difficult labor market circumstances such workers face. Alternatively, it is possible that the large gap relative to pre-recession expectations could reflect developments left out of our models. One possibility is that welfare reform in the late 1990s was a one-time boost to LFP that should not have been extrapolated into further LFP increases for low-education workers in the 2000s. Indeed, the largest gap today arises from female high school dropouts, the primary group affected by changes in the welfare laws. Another possibility is that the construction boom of the 2000s masked a longer-term deterioration of opportunities for low-education workers (Charles, Hurst, and Notowidigdo, 2014a, 2014b)—which could have also led our models to overestimate the trend in LFP.

Finally, in order to judge whether the level of actual LFP represents additional labor market slack over and above what is captured by the still somewhat elevated unemployment rate, one must ask whether LFP is low relative to expectations given the path of unemployment over the past few years. Accounting for those unemployment rates, plausible models for the trend LFP rate place the actual LFP rate between 0 and 0.8 percentage points below expectations. As noted earlier, there is ample reason for uncertainty about such estimates. However, our preferred estimate of 0.8 percentage points for the LFP gap would represent a nontrivial amount of additional labor market slack over and above that represented by the unemployment rate. In addition, our results suggest that compositional changes in the labor force may have reduced the natural rate of unemployment by up to 0.6 percentage points since 2000—a development not accounted for in prominent estimates of the natural rate. Such additional slack would suggest that monetary policy should remain more accommodative than would otherwise be the case.

## NOTES

[1]U.S. Bureau of Labor Statistics' (BLS) official LFP rate (available at http://data.bls.gov/timeseries/LNS11300000).

[2]The trend LFP rate is the LFP rate consistent with the contemporaneous composition of the work force and an economy growing at its potential.

[3]Our estimates of the actual and trend LFP rates reported throughout this article are computed from the U.S. Bureau of Labor Statistics' *Current Population Survey* (CPS). However, it should be noted that our actual LFP rate differs slightly from the official BLS LFP rate mentioned in the first paragraph of the article, probably because we do not use the composite estimation that the BLS does (we explain the difference in greater detail in note 16). We explicitly note where official BLS LFP data are used or referenced—as in figures 4, 5, 12, and A1 and related discussion.

[4]However, as we discuss in other parts of this article, there is evidence of changes in cohorts' LFP tendencies between youth and prime working age and also between prime working age and older ages.

[5]For the last age category, please note that the CPS does not distinguish age beyond 80 in some years.

[6]The unemployment gap is the gap between the actual unemployment rate and the Congressional Budget Office's (CBO) short-run NAIRU series. NAIRU stands for nonaccelerating inflation rate of unemployment and is one notion of the natural rate of unemployment, or the trend rate of unemployment. The natural rate of unemployment represents the unemployment rate that would prevail in an economy making full use of its productive resources. We further discuss this measure later in the text.

[7]See note 6.

[8]Trend (payroll) employment growth is the level of employment growth that is consistent with a flat unemployment rate. Employment growth above (below) trend will put downward (upward) pressure on the unemployment rate.

[9]A longer view of the LFP rate is available in the appendix's figure A1. These values are official numbers from the U.S. Bureau of Labor Statistics, *Current Population Survey*, from Haver Analytics.

[10]We plot in figure 2 the age-specific LFP rates for men and women in 2014. Because the data are from a cross section of a single year, they combine many birth cohorts rather than following a single cohort over the life cycle. We discuss this issue in more depth later. However, the overall shape of the life cycle would look similar if we followed birth cohorts over time or chose a base year other than 2014.

[11]While the female LFP rate remains about 10 percentage points below the male LFP rate, a number of studies suggest that women's labor force decisions—that is, how they respond to changes in wages, aggregate employment conditions, and public policies—now closely resemble those of men (Blau and Kahn, 2007; Heim, 2007; and Bishop, Heim, and Mihaly, 2009).

[12]See, for example, Juhn and Murphy (1997), Peracchi and Welch (1994), Autor and Duggan (2003), Blau (1998), and Blau and Kahn (2007). See also Juhn and Potter (2006) for a review. For details on DI and SSI programs, see www.ssa.gov/disability/.

[13]See, for example, Charles, Hurst, and Notowidigdo (2014a, 2014b), Autor (2010), and Acemoglu and Autor (2011), as well as the references therein.

[14]Data from the U.S. Bureau of Labor Statistics, *Current Population Survey*, from Haver Analytics.

[15]Ibid.

[16]As mentioned earlier, although the BLS uses the same basic CPS files to compute the official LFP rate series, our estimate of the aggregate LFP rate differs slightly. This difference may stem from the fact that in calculating the LFP rate, we do not use the composite estimation that the BLS does; this estimation exploits the CPS's rotation sample design (with households in the survey for four months, then out for eight months, and finally in again for four months). The sample rotation scheme results in a positive correlation between CPS estimates from different months, improving measures of change over time. The CPS composite estimate for a given labor force statistic (for example, the number of people unemployed or employed) is based on a weighted average of two estimates for the same statistic: 1) the CPS estimate and 2) the previous month's composite estimate plus an estimate of change since the previous month. In addition, the composite estimate also incorporates an adjustment to partially correct for bias associated with time in the sample (by assigning higher weights to data from households completing their first and fifth interviews in the month).

[17]The data series for these controls are plotted in the appendix's figure A2. The series for marital status and the presence of a young child (under five years old) are computed from the CPS data. The ratio of youth to adult wages is computed at the state level from the CPS microdata using average hourly wages of those paid at an hourly rate. (Youth is defined as 16–24 year olds and adult as 25–54 year olds.)

[18]The CBO short-run NAIRU series—which accounts for temporary factors, such as unemployment insurance extensions—is from Haver Analytics. See also note 6 for further details.

[19]The state unemployment rates are from Haver Analytics.

[20]See www.dol.gov/whd/state/stateminwagehis.htm.

[21]One issue with these alternative labor market measures is, as far as we know, there are no standard estimates of their long-run trends. For the state-level unemployment rate, we adjust the national natural rate of unemployment for the deviation of the state unemployment rate relative to the national unemployment rate averaged over the estimation sample period. Specifically, the adjusted state-level unemployment gap is $u_{s,t} - u_t^* - (\bar{u}_s - \bar{u})$ and its lags. Note that state is the most detailed geographic unit that includes all CPS respondents. For the unemployment spell duration measure, we include the median spell (in additional to the rate) of national unemployment. We use median instead of mean spell because the CPS recorded unemployment spell duration up to two years through 2011 and up to five years thereafter, which causes a discrete change in mean duration but leaves median duration intact. To isolate a cyclical component in duration, we used deviations of median spells from the sample time period's mean.

[22]Both wage variables are measured as deviations from the sample time period's means.

[23]Ideally, we would also condition life expectancy on education, since mortality has varied over time by education levels (Meara, Richards, and Cutler, 2008). However, we have been unable to find a time series on education-specific life expectancy with a high enough frequency.

[24]Because of the negligible sample sizes, we do not estimate the model for those with a postcollege degree aged 16–24. This leaves us with (3 × 2 × 5) – 2 = 28 groups. Additionally, we exclude from the data small numbers that are high school graduates younger than 17, high school graduates with some college younger than 19, and college graduates younger than 21.

[25]The parameters $\lambda_{se}$, $\gamma_{se}$, and $\delta_{se}$ are the regression coefficients on the *w*, *x*, and *z* variables, respectively.

[26]We also allow for a trend break in 1992 to account for the redesign of the education question in the CPS (see note 1 of box 1, p. 114) and to extrapolate the trend based on the more recent and relevant time period.

[27]The parameter $\theta_{se}$ is the regression coefficient on the *t* variable.

[28]Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*.

[29]The model includes quarterly dummies to adjust for seasonal patterns in work. Later on, we also discuss the results when the model is run through other time periods, including from the beginning of 1982 through the third quarter of 2014.

[30]Keep in mind that each estimate is based on a single birth year and further stratified by sex and education level. Therefore, the series plotted in panels A, B, and C of figure 8 are noisy. For our purposes, the noise washes out once we aggregate many series together to form our national trend LFP rate.

[31]However, the cohort effects on the later career work activity of women aged 55–79 (not shown in figure 8) have generally been rising over time.

[32]Recall that we stop our estimation sample in 2007 and forecast trend LFP for 2008 and onward. For forecasts for 2008–13, we use the actual population data.

[33]See Bell and Miller (2005).

[34]The actual LFP rate fell by 0.5 percentage points in the fourth quarter of 2013. As of third quarter of 2013, the actual LFP rate was only 0.3 percentage points below where we would have expected given economic conditions.

[35]The derivation of the equation uses the fact that $\sum_d p_{t-1}\Delta f_{dt} = p_{t-1}\sum_d \Delta f_{dt} = 0$. The second equality follows because $f_{dt}$ is a share and always sums to 1 at a given *t*.

[36]We examine the population aged 25 and older in order to exclude younger individuals who might not have completed their education. Of course, many individuals continue their education beyond age 25. See, for example, figure 5 in Aaronson and Sullivan (2001).

[37]The rising LFP rate of high school dropouts over this period is mainly due to the increasing LFP of prime-working-age women. Many researchers have studied the implications of policy changes, such as welfare reform and the expansion of the earned income tax credit (EITC), on the LFP decisions of female low-skill workers (Eissa and Liebman, 1996; Meyer and Rosenbaum, 2001; Moffitt, 2003; and Eissa and Hoynes, 2006).

[38]Specifically, this model includes single-year age dummies, single-year birth cohort dummies, dummies for the three baseline age-groups (16–24, 25–54, 55–79) interacted with the state unemployment gaps as well as the race and quarter dummies, and subsets of age group dummies interacted with group-specific controls (for instance, a dummy for ages 16–24 interacted with minimum wage and with youth-to-adult wage ratio and a dummy for ages 25–54 interacted with marital status and the presence of children under five years old).

[39]Technically, we use sampling weights in our logit regression analysis. Consequently, the standard likelihood ratio (LR) test or the Akaike information criterion (AIC)—methods of measuring the relative quality of a statistical model—cannot be used. Instead, we apply these tests to an unweighted version of our baseline model, which turns out to give nearly identical estimates as the weighted version. We find that for all ten sex-education-level groups, the LR test can strongly reject the restrictions imposed by the pooled model (with *p*-value < 0.001). Moreover, the AIC also favors the less restricted three-age-group (baseline) model across all sex-education-level groups.

[40]Adjusting the natural rate of unemployment for changes in the educational distribution of the labor force is somewhat controversial. Summers (1986) argues that such adjustments imply counterfactually high unemployment rates in earlier years. Shimer (1999) builds a model in which workers' relative levels of education signal ability to employers, but average absolute levels of education do not affect unemployment. However, we see fairly modest empirical support for education signaling models. Altonji and Pierret (2001) show that employers use education level as a proxy for unobserved productivity among job applicants but that this signaling effect fades once firms learn new information about the productivity of their hires. Lange (2007) builds a model to quantify the speed of employer learning about new workers and shows that, under his preferred specification, only 10 percent of the workers' return to schooling can be ascribed to education signaling. Clark and Martorell (2014) provide direct evidence that education signaling may not matter to wages in the case of high school diplomas. Shimer (1999) also notes that endogenous choice of schooling levels might bias upward the effects of education on unemployment if more-able people choose to get more education. That said, research on the effects of education on wages using plausible instrumental variables or twins-based designs does not produce estimates notably below those obtained from ordinary least squares; see, for example, Card (1999).

Our preferred interpretation of the impact of schooling on unemployment is that increased education has indeed been pushing down unemployment for many decades, but that its effects have been offset by other factors. However, given that there is uncertainty over whether adjustments for education are warranted, we note that changes in the age distribution of the labor force alone lower the natural rate of unemployment by 0.34 and 0.16 percentage points relative to what it would have been in the third quarter of 2014 if the age composition of trend LFP had remained the same as in 2000 and 2007, respectively. Similarly, changes in the distribution of educational attainment of the labor force alone lower the natural rate of unemployment by 0.40 and 0.23 percentage points relative to what it would have been if the composition of trend LFP had remained the same as in 2000 and 2007, respectively. Changes in the gender composition have no impact. Together, changes in the age distribution, gender composition, and the distribution of educational attainment reduce the natural rate of unemployment by 0.61 and 0.32 percentage points relative to what it would have been if the composition of trend LFP had remained the same as in 2000 and 2007, respectively.

[41]As mentioned earlier, the CBO's short-run NAIRU accounts for temporary factors, such as unemployment insurance extensions, that boosted the natural rate after 2007. The long-run NAIRU does not include these transitory factors.

[42]Specifically, we adjust this hypothetical natural rate of unemployment by $u_t^{hypo} + \left(u_t^{CBO\_LR} - 5.0\right)$ for $t \geq 2008$.

[43]The trend employment-to-population ratio $\frac{\hat{e}}{pop} = \hat{p} \times (1 - \hat{u})$, where $\hat{p}$ is trend LFP and $\hat{u}$ is the natural rate of unemployment.

[44]For the last constructed measure, note that payroll employment is the employment reported in the BLS's *Current Employment Statistics* survey, which is also referred to as the payroll or establishment survey. Household employment is from the CPS. For further details on these two different measures of employment, see www.bls.gov/web/empsit/ces_cps_trends.pdf.

[45]Both the historical data and projections for the civilian noninstitutional population aged 16 and older are from the U.S. Census Bureau. We use the U.S. Census Bureau's national Quarterly Intercensal Noninstitutional Civilian Population files (1982:Q1–1990:Q1) and the Monthly Postcensal Noninstitutional Civilian Population estimates (1990:Q2–2013:Q4). The U.S. Census Bureau's 2013–20 population projections are from the 2012 National Population Projections, which were released on December 12, 2012. Historical data are found at www.census.gov/popest/data/historical/index.html and the projections at www.census.gov/population/projections/data/national/2012/downloadablefiles.html. Discontinuities between the two series are smoothed. We also smooth the data to adjust for revisions produced by decennial censuses. We adjust for the seasonal pattern in population shares by using a four-quarter moving average. We then use the Hodrick–Prescott (HP) filter to isolate a trend component. To avoid the standard end-of-sample problem with the HP filter and because the U.S. Census Bureau's projection of trend population is superior to a statistical estimate from an HP filter, we replace the HP-filtered trend with the U.S. Census Bureau's projections after 2015.

[46]As with population growth, we use the HP filter to estimate the trend of the ratio of payroll to household survey employment.

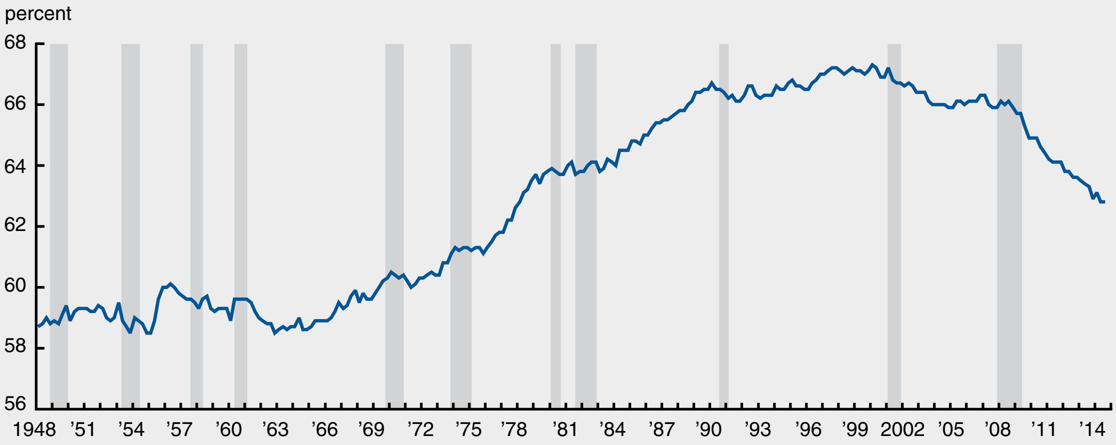
[47]The increase and subsequent decline in the ratio of payroll to household survey employment is evident even when using the payroll-concept-adjusted household employment series. This series is a research series created by the BLS to make the household employment series more comparable to the payroll employment series (see note 44); for details, see U.S. Bureau of Labor Statistics (2012).

[48]The final trend employment growth series is smoothed using a four-quarter moving average.

APPENDIX

**FIGURE A1**

**Labor force participation rate, 1948–2014**




Notes: The figure plots official quarterly data over the period 1948:Q1–2014:Q3. The shaded bars indicate recessions as defined by the National Bureau of Economic Research.
Source: U.S. Bureau of Labor Statistics, *Current Population Survey*, from Haver Analytics.

**FIGURE A2**

**Additional employment-related and demographic data series used as controls for estimating the statistical models, 1982–2020**

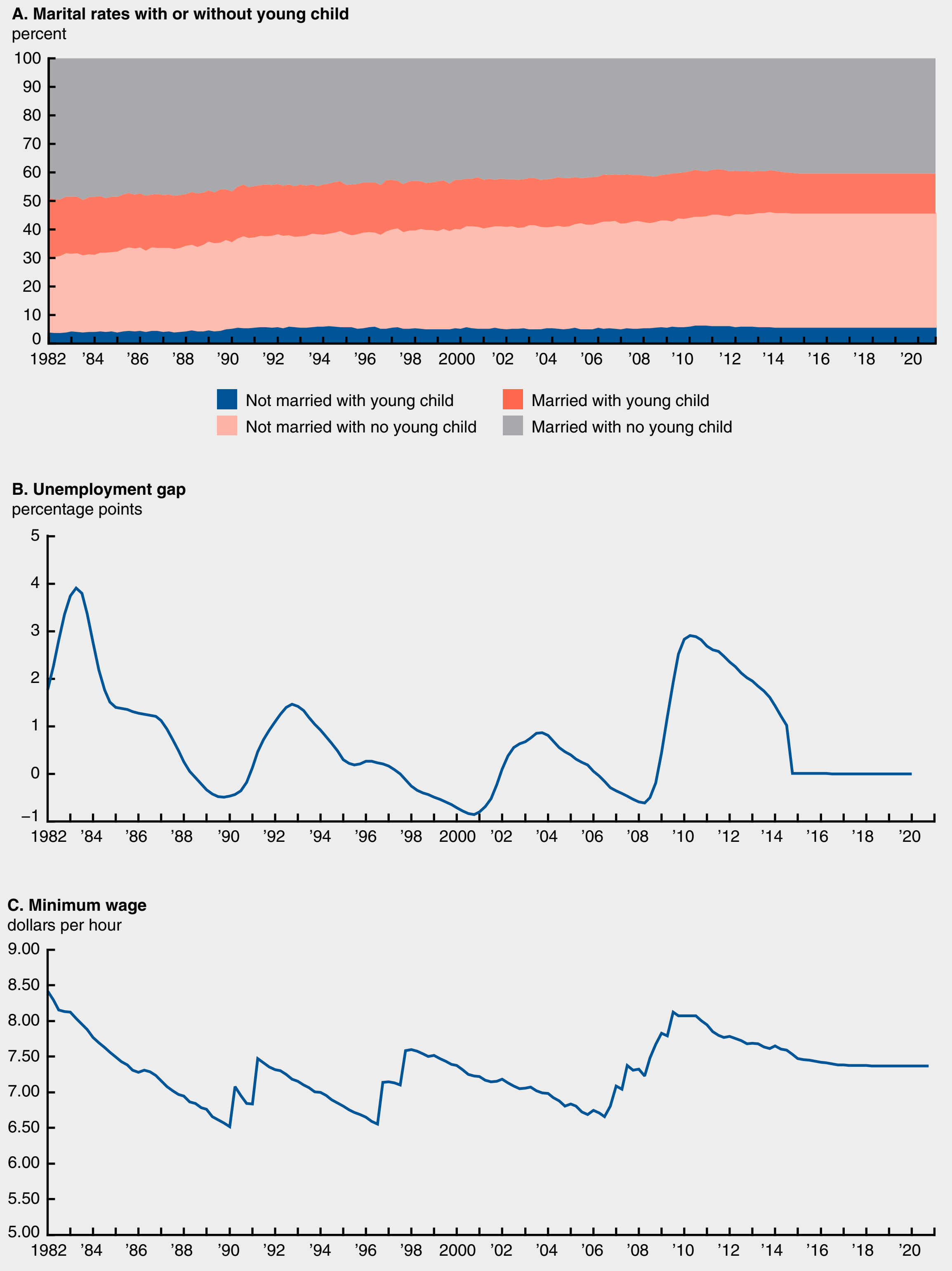
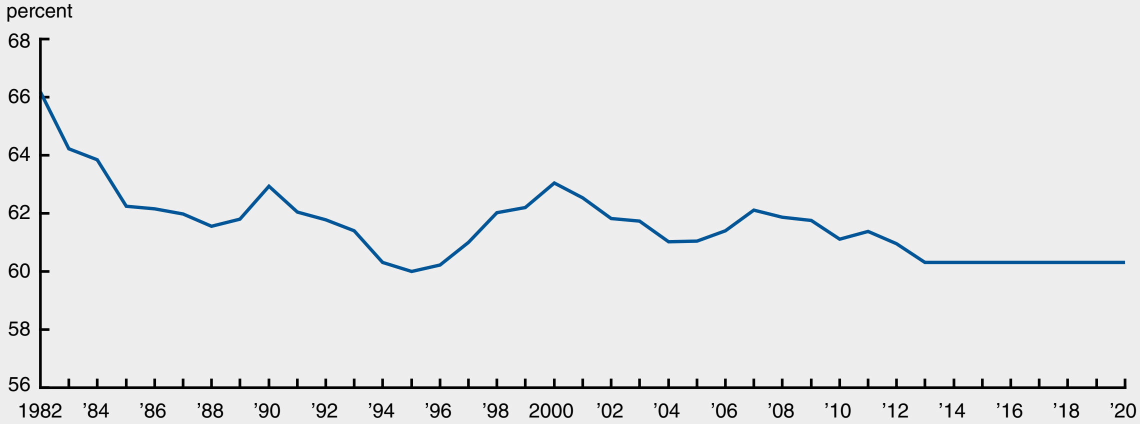


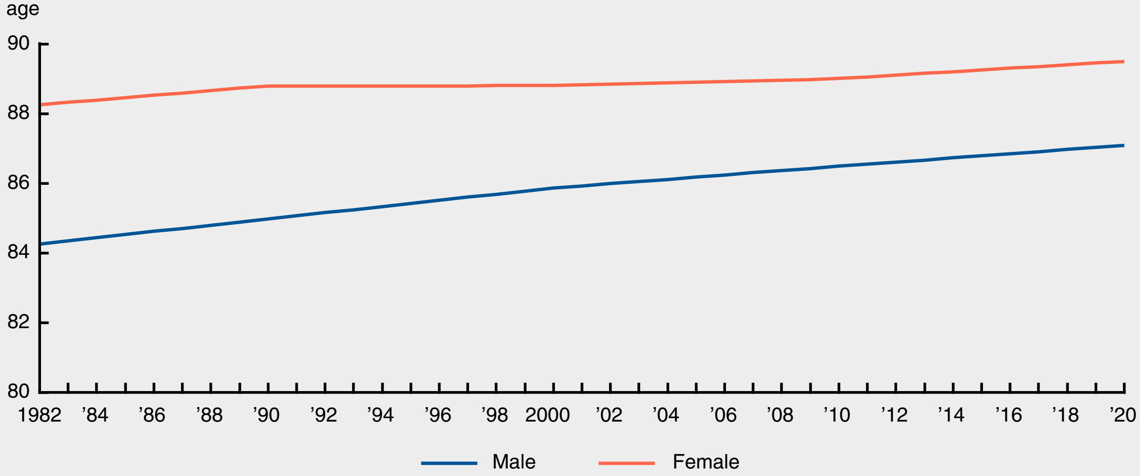
**FIGURE A2 (CONTINUED)**

**Additional employment-related and demographic data series used as controls for estimating the statistical models, 1982–2020**

**D. Youth-to-adult wage ratio**




**E. Life expectancy**




Notes: For panel A, a young child is a child under age five. For panel B, the state-level unemployment gap (that is, the difference between the actual unemployment rate and the CBO’s short-run natural rate of unemployment explained in the text) is a four-quarter moving average with no lag. For panel C, the state minimum wages are deflated by the BLS’s Consumer Price Index for All Urban Consumers and then averaged. For panel D, youth are aged 16–24 and adults are aged 25–54. For panel E, the life expectancy plotted is that averaged for all ages 16 and older. All the panels plot quarterly data over the period 1982:Q1–2020:Q4. Methods for obtaining projections after the current quarter are described in the text.

Sources: Authors’ calculations based on data from the U.S. Bureau of Labor Statistics (BLS), *Current Population Survey*, from Haver Analytics; Congressional Budget Office (CBO) from Haver Analytics; U.S. Bureau of Labor Statistics, January issues of the *Monthly Labor Review*; U.S. Department of Labor (www.dol.gov/whd/state/stateminwagehis.htm); and Bell and Miller (2005).

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