Are Less Skilled Women Crowding Out Labor Market Participation Among Less Skilled Men?

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The recent declines in labor supply among less skilled men have been clearly documented in the first paper in this volume (Holzer and Offer, 2001). That paper and others also clearly document the increase in labor force participation among less skilled women. For instance, Blank and Schmidt (2001) indicate that labor force participation among women without a high school degree increased by six percentage points between 1989 and 1999; single mothers with children increased their labor force participation by over 10 percentage points during this decade. This paper investigates whether there is any evidence of a causal relationship between increases in female labor supply and declines in male labor supply among less skilled workers.

1. Changes in Female Labor Supply

There is a substantial literature on female labor supply changes over the past decade, much of it driven by interest in the work-oriented welfare reforms that were enacted in the 1990s. A number of papers investigate various measures of female labor supply and attempt to determine how much of these changes are explainable by the booming economy of the 1990s and how much are due to policy changes (Moffitt, 1999; Schoeni and Blank, 2000; or Grogger, forthcoming).

There are several policy changes on which these papers focus. First, they investigate the effects of state waivers, which allowed states in the early 1990s to petition the federal government to operate welfare programs whose rules varied from the Aid to Families with Dependent Children (AFDC) program, the primary cash assistance program to single mothers. Most of these waivers allowed states to run stronger welfare-

to-work programs, often including time limits and strong sanctions for noncompliance with work requirements.

Second, these papers investigate the effect of the 1996 welfare reform legislation. In 1996, the Federal government abolished AFDC and gave states much more discretion over the design of cash assistance programs. Federal dollars that had been funding AFDC were converted to the Temporary Assistance to Needy Families (TANF) block grant. Virtually all states took advantage of this change to enact substantially revised TANF-funded welfare programs.[1] While these new welfare programs varied substantially across states, they typically involved strong efforts to move welfare recipients into work, provided both incentives to work and penalties for not cooperating, implemented national and state time limits, and tried to discourage women from ever entering welfare through various diversion efforts.

These welfare program changes increased the costs of not working and in many cases also increased the returns to work. Several other policy changes have further increased less skilled women's incentives to work. Substantial increases in the minimum wage over the 1990s improved wage rates for low wage workers. In inflation-adjusted dollars, minimum wages were 10.8 percent higher in 2000 than in 1989. These minimum wage increases were supplemented by real wage growth stimulated by the strong economy; virtually all less skilled workers increased their real wages over the late 1990s (Blank and Schmidt, 2001).

In the early 1990s, Congress also enacted substantial increases in the Earned Income Tax Credit (EITC), a program that supplements wages for low wage workers in

[1] For a summary of these legislative changes, see Blank and Ellwood (2002, forthcoming); for a summary of the estimated impacts see Blank (forthcoming).

low income families through the tax system. The combined increases in minimum wage and EITC helped increase the returns to work and have been clearly linked with increases in the labor supply of single mothers (Eissa and Leibman, 1996; Meyer and Rosenbaum, 2001; Ellwood, 2001).[2]

A final set of policy changes further increased the returns to work for single mothers by reducing work-related expenses through expansions in child care subsidies and access to public health insurance. Federal dollars for child care subsidies to women on welfare rose from $0.6 billion to $6.6 billion between 1988 and 1999 (Blank and Ellwood, Figure 1).[3] Publicly provided health insurance (Medicaid) has also been made more broadly available to women who leave welfare for work, and eligibility for Medicaid is now automatic for all children in very low income families.[4] In the past, families needed to be explicitly eligible for certain assistance programs in order to receive public health insurance; because so few low wage jobs offer health insurance coverage, leaving welfare and going to work often meant losing health insurance.

All of these programmatic changes should have substantially increased the returns to low wage work and reduced access to benefits for nonworkers, especially among single mothers with children. Of course, the booming economy of the late 1990s also made jobs more available and raised wages. The U.S. unemployment rate remained at or

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[2] In contrast, Eissa and Hoynes (1998) indicate that these changes had small negative effects on the incentives for married women to work.

[3] This is a minimal estimate of the growth in child care subsidies. It is based on federal programs that provide direct subsidies to welfare recipients. It does not include state child care subsidy dollars which have also expanded. And it does not include expansions in the Dependent Care Tax Credit that benefits families whose income is high enough to owe taxes. See Loprest, et. al. (2000) for more details on child care subsidy changes.

[4] All children below age six in families whose income is less than 133 percent of the federal poverty line have access to Medicaid; 23 states use higher income cutoffs. All children up to age 19 in families whose income is less than 100 percent of the poverty line are also covered by Medicaid; 26 states use higher income cutoffs. For more details on these Medicaid expansions, see Gruber (2002, forthcoming).

below five percent from April 1997 until October 2001. Much of the recent research literature focusing on female labor supply changes has attempted to measure the impact of these economic and policy changes.

The relative importance of these different factors varies across studies and policies. For instance, both Moffitt (1999) and Schoeni and Blank (2000) conclude that waivers had a significant and positive impact on work behavior among less-skilled women. But Schoeni and Blank find few effects of TANF implementation on labor market participation; once they control for state economic environment the policy changes appear to matter little post-1996. Unfortunately, there are methodological problems with trying to estimate the impact of all of these policy changes at the same time, so we have few direct comparisons of the simultaneous role of the economy, the EITC and minimum wage expansions, welfare changes, and other related policies.[5] A good summary of the research, however, is that both the booming economy of the 1990s and a host of related policy changes appear to have increased female labor supply. Blank (2000) argues that there were synergies between the various policy changes and the high rate of economic growth, so that all of these effects reinforced each other.

2. Possible Links Between Male and Female Labor Supply

Of course, the same factors that produced changes in female labor supply may directly impact male labor supply as well. Certainly the strong economy and rising real wage rates among less skilled workers should provide incentives for both low-wage men and women to enter the labor force at a higher rate. For instance, inflation-adjusted

wages among male high school dropouts rose 3.5 percent between 1994 and 1999 (Blank and Schmidt, 2001). Similarly, the effects of policy changes (like expanded health insurance or child care subsidies) that reduce the costs of work for parents should also increase male labor supply among less skilled men who are single fathers or married men with children.

Increases in disposable income among low income working married couples could result in intra-family labor market substitution, however. Expansions in work opportunities and wage rates for both men and women could increase family income and (through an income effect) provide married men with an opportunity to work somewhat less. In most cases, however, we assume that it is the wife who has the strongest response to increases in family income by substituting away from labor supply. This substitution makes the most sense for the lowest-wage worker in the family and husbands' wage rates are typically higher than are wives'. Hence declines in male labor supply would occur only if both parents earn equally low wages and if increases in the returns to work result in income effects on family labor supply. Such a response would be surprising, however, since past evidence suggests that married men respond positively to increased returns to work (Blundell and MaCurdy, 1999).

In general, the welfare policy changes should have had relatively few direct effects on male labor supply. About 10 percent of welfare recipients in the 1990s were male-headed families - either single male parents or married couples who were eligible for cash benefits. The AFDC program, like state TANF programs after 1996, established explicit provisions under which married couples could receive welfare. In most cases,

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[5] For instance, the methodologies used to measure the impact of the economy and of welfare reforms typically use year fixed effects, which makes it impossible to separately identify the effects of minimum

however, it was harder for married couples to get cash assistance and hence very few of them participated. Among all low skilled women (high school degree of less), xx percent reported receiving welfare income in 1995; among low skilled men, only yy percent reported receiving welfare <get figures>. This suggests to us that the welfare reforms of 1996 and earlier should have had few direct effects on men.

Similarly, neither welfare reform nor other policies designed to encourage work or improve the returns to work among low wage parents (such as the EITC expansions or the expansion in child care subsidies) should have had much of a direct effect on single men or married men without children. As it turns out, a disproportionately high share of less skilled men are unmarried or in families without minor children. In the March 1996 CPS, 44 percent of all less skilled women are in families that do not include unmarried children age 18 or below. But only 56 percent of men are in such families.

If low-wage men are sharing income with girlfriends or family members, however, all of these policies could indirectly affect male labor supply, even among childless and unmarried men, as increases in work effort and wages among the women influence male labor market decisions. This can happen in two ways. First, there may be an indirect income effect on these men. If girlfriends or family members are working more and earning more, the men may feel less of a need to provide them with income. This could reduce male labor supply.

Second, in situations where female labor supply expands rapidly there may be limits to the number of available jobs. Given the booming economy, we expect that labor demand was expanding through most of the 1990s. If the only thing happening in the 1990s was an increase in labor demand, then the rightward shift in a demand curve would

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wage changes and EITC changes that were enacted everywhere at the same time.

pull up wages and push out employment. One would expect these benefits to be shared by both male and female workers. But what if the increase in demand was entirely satisfied by an expansion in female labor supply? The welfare-to-work policies enacted in the 1990s may well have pushed out labor supply among female workers. In the language of economics, the expansion in female labor supply might have been due to both a shift in the demand curve (which results in increased employment as one moves up the supply curve to a higher level of demand), as well as shifts in the supply curve of female labor, due to the strong enforcement of welfare-to-work efforts that pushed more women into the work force. It is possible that the outward expansion of the female labor supply curve more than accommodated the rise in labor demand, perhaps reducing the number of jobs available for men.

This would be a surprising outcome for two reasons. First, the facts of the late 1990s do not support this simple story. Real wages rose among *both* men and women, and unemployment rates remained low among *both* men and women. A demand substitution story, in which women are hired and men are fired, should result in falling wages and rising unemployment among less skilled men and that did not happen. Indeed, the low overall unemployment rates and constant employer complaints about not having enough labor throughout the late 1990s suggest that there were jobs available for both less-skilled men and women.[6]

Second, most evidence suggests that men and women are in somewhat different labor markets. The degree of occupational segregation among less skilled workers is high, with men more likely to be in manufacturing jobs and women more likely to be in

[6] Of course, it is always possible that unemployment among men would have fallen faster and/or wages risen faster in the absence of such strong increases in female labor supply.

service sector jobs. Under these circumstances, even a rapid expansion in female labor supply should not reduce labor demand one-for-one in traditionally male occupations.[7]

All of this suggests that something more than a simple displacement story had to be occurring in the late 1990s. (Although, of course, in some areas with very rapid increases in female labor supply displacement could have occurred, even if it is not the primary cause of male labor supply changes in aggregate.) Perhaps rising income among mothers, sisters, and girlfriends resulted in reduced incentives for men to work, as suggested above. Perhaps there were compositional changes in *which* men were working over the 1990s. Perhaps the booming economy created job opportunities in the underground economy that pulled men away from mainstream employment. Perhaps employers came to prefer women over men, as the changing mix of low-skilled jobs made the "soft skills" at which women excel more important to employers. These latter explanations are explored in other papers in this volume.

In the remainder of this paper, we ask whether there is any evidence that rising female labor supply was associated with declines in male labor supply - either through displacement or through income effects. We refer to this as the "substitution hypothesis." Certainly the overt evidence that female labor force participation boomed as male labor force participation fell lends at least some credibility to this hypothesis and makes it worth pursuing in the data.

Seeking to identify substitution effects between two closely-related groups in the labor market is difficult. In particular, most measures are troubled by simultaneity problems. By this, we mean that underlying economic forces drive labor force

[7] In the 2000 ORG data, the occupational gender segregation index for workers with more than a high school degree is 0.48, while it is 0.57 for those with a high school degree or less.

participation responses of both men and women in the same direction. This means there will be a strong positive correlation between male and female labor supply, which implies nothing about the causal effect of changes in female labor supply on male labor supply. One must somehow get rid of this positive bias in a credible way in order to claim that one has tested the substitution hypothesis. Because supply and demand move simultaneously and because the market for male and female labor overlaps heavily, it is very difficult to separate out a potential "substitution" effect from other possible correlations and causes.

While we know of no other literature that focuses generally on male-female labor market substitution in the booming labor market of the 1990s, others have tried to measure the potential displacement effects of increases in labor force participation among welfare recipients. Bartik (2000) summarizes this literature and contributes a variety of estimates using different theoretical approaches to measuring displacement effects. He concludes that there is little evidence that welfare reform alone had large impacts on the overall labor market, but it may have had larger effects among less skilled workers. His estimates suggest the largest effects might occur among other less skilled women rather than among less skilled men.

Our primary way of dealing with simultaneity in the determination of male and female labor supply is to try several different types of estimation procedures, all with some admitted faults, to see if any of them result in evidence of substitution. In the next section (section 3), we look at the comparative changes in the labor supply of less skilled male and female workers within defined labor markets, trying to control for economic, demographic, and institutional factors. In section 4 we investigate the effects of welfare

policy changes on male labor supply. If these are primarily affecting work behavior among less skilled women and affecting men only indirectly through their impact on female labor supply, then any measured negative effect of these policy changes on male labor supply indicates some sort of substitution effect. (In the language of econometrics, we use these policy changes as an instrument for female labor supply.) In the final empirical section (section 5), we focus on transitions into employment and full-time employment, to see if male transition rates are affected by female transition rates. Transitions may be an important margin of adjustment if crowding out happens at the hiring stage in an expanding economy.

3. Investigating Changes in Male and Female Labor Supply Within Identifiable Markets

In this section we focus on changes in male and female labor supply within defined labor markets. The labor markets we use are Metropolitan Statistical Areas (MSAs). MSAs include all cities with over 50,000 inhabitants and the county or counties in which they are located and with which they are economically and socially integrated. We utilize 208 MSAs in the analysis in this section. These are the MSAs that were defined in both 1989 and 2000 within the data.[8]

The data come from the Outgoing Rotation Groups (ORGs) of the Current Population Survey (CPS). The ORG data provides wage and employment information on

[8] We are not concerned with changes in MSA boundaries between these two years. Our interest is the metropolitan labor market. As MSAs grow, their relevant labor market necessarily expands.

a monthly national sample of respondents.[9] Aggregating the 12 monthly samples collected in each year provides enough observations in each of these 208 MSAs to be able to look at skill groups by gender. We primarily focus on men and women with less than a high school degree or exactly a high school degree, to who m we refer as less skilled workers. Our unit of observation is the MSA average within a specific gender and education grouping. Sample sizes within these cells range from 2275 observations on less skilled men in the largest MSA in 2000 to 19 in the smallest MSA.

Figure 1 shows the changes in male labor supply within our sample over time between 1989 and 2000. Male labor force participation among those with only a high school degree declines from 81 percent to 76 percent over this period; among those with less than a high school degree, labor force participation declines from 56 percent to 53 percent in 1994, but then goes back up to 55 percent by 2000.

Since changes in male labor supply can occur along both the intensive as well as the extensive margin, Figure 2 shows changes in mean hours among those who work. It is difficult to interpret trends in these data since the hours definitions change between 1993 and 1994. There is little evidence, however, of major changes in hours over time either before or after this breakpoint.

The simple correlation between changes in male labor supply and changes in female labor supply between 1989 and 2000 within these 208 MSAs is provided in Table 1, row 1. Column 1 shows the correlations in male-female labor supply changes among all workers and columns 2 and 3 show the correlations among male-female labor supply changes among those with less than a high school degree and among those with exactly a

[9] The monthly CPS sample collects employment information. The ORG sample is composed of the outgoing rotation groups (one-fourth of the CPS sample); these individuals are also asked wage

high school degree. There is a relatively strong and positive correlation in all three columns, meaning that areas with larger increases in male labor supply were also areas with strong increases in female labor supply for all skill levels. This is not surprising, and almost surely indicates the dominance of economic factors in a labor market that affect both men and women in similar ways.

Column 2 shows correlations between changes in hours of work among male workers and female labor force participation changes. These correlations are negative but are essentially zero, indicating little correlation between hours worked and female labor supply changes.

The third and fourth rows of Table 1 provide a comparison between changes in male labor supply in the 20 MSAs where labor force participation among less skilled women (those with a high school degree or less) increased the fastest over this decade (third row), and changes in male labor supply in the 20 MSAs where changes in less skilled female labor supply increased the least over the decade (fourth row). The fifth row provides a difference-in-difference estimator of the changes in male labor supply over time and between areas with high and low changes in female labor force participation.

A simple displacement story would suggest that male labor supply should fall more (or rise less) in areas of fast-growing female labor supply versus areas of slow-growing female labor supply. The evidence suggests just the opposite. Figure 3 reinforces this by plotting labor supply changes among low skilled men in the 20 fastest versus the 20 slowest growth areas in terms of female labor supply. It is clear that male labor supply is falling much faster (by 7 percentage points) in the areas with lower

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information.

increases in low skilled female labor supply, while it hardly changes at all in the areas with high rates of female labor supply growth.

By itself, this evidence proves little. Without controls for the economy, these correlations are likely to be driven by common economic effects within metropolitan areas. But these simple estimates do establish that there is no prima faciae evidence of substitution in these data.

Table 2 provides a set of somewhat more controlled regressions using a panel of MSA-specific data for the 12 years from 1989 to 2000. The table indicates what happens to the correlation between changes in male and female labor supply once other variables are controlled for. The dependent variable in Part I is the change in male labor force participation by MSA and year. Columns (1) to (3) utilize data on all men; columns (4) to (6) utilize data on all less-skilled men (those with a high school degree or less); while columns (7) to (9) utilize data only on men with less than a high school degree. Row one shows the coefficient on the change in female labor force participation across three increasingly rich specifications for each skill group. With 208 MSAs and 12 years there are 2496 observations in each regression.[10]

Columns (1), (4), and (7) show the coefficient on change in female labor supply, with MSA and year fixed effects included. Columns (2), (5), and (8) also control for the unemployment rate (both current and lagged one period), indicating the effect of controlling for differences in economic environment between areas and over time. Columns (3), (6), and (9) add controls for a variety of additional demographic and labor market variables, including percent black and percent Hispanic in the MSA labor market,

and percent employed in manufacturing, percent in nonprofessional services, percent in professional services, percent in public administration, and an index of gender industry segregation. This industry segregation index is calculated in each MSA in each year among all workers.[11] We might expect that labor markets with greater gender segregation would show smaller substitution effects.

The results in Table 2 continue to show few indications of substitution in this data set (which should generate a negative relationship between female labor supply changes and changes in male labor force involvement if we have adequately controlled for economic conditions.) In all cases, coefficients showing the effect of changes in female labor supply on changes in male labor supply are positive and significant for all specifications and skill levels, even once economic environment, demographic differences and industry structure are controlled for. The inclusion of MSA and year fixed effects in all regressions further controls for any common factors within an MSA over time or across all MSAs in a given year. Columns (3), (6), and (9) offer a relatively rich specification that should control for many of the things that cause a positive bias in the relationship between male and female labor supply.[12] Yet, even in these columns the coefficient on female labor supply is positive and significant for all education groups.

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[10] Table 2 regresses changes in male labor supply on changes in female labor force participation. We have also regressed levels of male labor supply on changes in female labor force participation and a lagged dependent variable and found similar results.

[11] This is known as a Duncan index and computed as $(.05)*?_i(M_i - F_i)$ for each MSA in each year. M_i is the proportion of the male labor force employed in industry i and F_i is the proportion of the female labor force employed in industry i. (We calculate the index across 23 major industries in each MSA.) The measure indicates the proportion of women (or men) who would have to change industries for the industry distribution of men and women to be the same. Values range from 0 to 1, with 1 indicating total segregation.

[12] Since the dependent variables in these regressions is the change in participation rates, fixed effects are removing systematic differences in LFP trends.

There is little sign that the relationship for the least skilled group of men (columns 7 through 9) is significantly different than that among all men.

Part II uses changes in mean hours of work among men who work as the dependent variable, testing whether the degree of male labor market involvement is affected by changes in female labor force participation. The coefficients show the effect of changes in female labor supply on hours worked among men across the same three specifications shown in Part I and for the same skill groups. In this case, the coefficients on change in female labor force participation are small and insignificant in all specifications and for all skill groups.

The results in this section should not be overinterpreted. Because of the simultaneous nature of labor supply responses among men and women in the same labor market, one might expect to find little evidence of substitution. The main message of this section is clear, however: There is no first-order evidence in these data of substantial substitution between men and women in the labor market over the 1990s. While there may be small effects that can be found with more sophisticated estimation techniques, these effects are likely to be second-order. Any large substitution effect should show up in the specifications in columns (3), (6) and (9) of Table 2, which control for a relatively rich set of environmental variables.

4. Using Policy Changes as a Potential Instrument for Female Labor Force Changes

A problem with all of the above calculations is that both male and female labor supply are simultaneously affected by the economic environment. This should lead to a strong positive association between the two variables which cannot be interpreted

causally and which will bias upward any estimated coefficient of the effect of female labor force participation on male labor force participation. The classic response in such a situation is to look for an instrument, that is, a variable that is correlated with female labor supply and not with male labor supply. One could then use this variable as a proxy for female labor supply changes and measure only the causal impact of female labor supply on male labor supply, eliminating the positive bias due to other variables that simultaneously affects male and female labor supply together.

In most cases finding an instrument can be difficult; in this case there is an obvious choice: the policies enacted during the 1990s that were designed to move welfare recipients - predominantly single mothers -- into work. In order for these policies to serve as an effective instrument, they must have had a directly effect only on female labor force participation. Their effect on male labor force participation must come solely through the influence of female labor force changes.

As we noted above, there are some problems with this assumption. A small share of welfare recipients are male-headed families. However, as we also noted, very few less skilled men receive welfare in any given year, while a substantial share of less skilled women are welfare recipients. This difference suggests to us that policy changes in welfare programs are likely to have few direct effects on male labor supply, and hence may function effectively as an instrument that is strongly correlated with changes in the labor supply of less skilled women.[13]

We use the approach first utilized in Schoeni and Blank (2000) to estimate the effects of welfare program parameter changes on female labor supply measures, except

we use male labor supply measures as our dependent variable. Schoeni and Blank estimate the effects of waivers, enacted in various states between 1992 and 1996, on female labor supply. They also estimate the effects of the implementation of TANF-funded programs in the mid-1990s. Like them, we utilize data by state, by year, by skill level (less than high school, high school, and more than high school) and by age group (16-25, 26-44, and 45-64).[14] We calculate male labor force variables for each age-year-state-education cell and use these as the dependent variables in a panel data set.

Our estimation equation is

$$(1)\ Y_{aest} = Waiver_{st} \bullet Educ \bullet \beta_{eWaiver} + TANF_{st} \bullet Educ \bullet \beta_{eTANF} + \gamma_s + \gamma_t + trend \bullet \gamma_s + Z_{aest} \bullet \beta_z + \varepsilon_{aest}$$

The variables are defined for women in age group *a* and education group *e* who live in state *s* in calendar year *t* as follows:

Y: The outcome variable (either labor force participation or hours among male workers.)

Waiver: An indicator variable that takes the value of one if the state in which the man lived had a major waiver in effect; the indicator is turned off when TANF is implemented in the state.

TANF: An indicator variable that takes the value of one if TANF was in effect in the state in which the man lived; the TANF implementation date varies across states between 1996 and 1997.

Educ: Indicator variables for education groups: Less than 12 years; exactly 12 years; and 13 years or more.

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[13] For a more extensive and nuanced discussion of the relationship between welfare reform and male labor supply (including specific attention to Black men), see Lewis (2002). Bartik (2000) also estimates some models of this sort.

[14] We use three age categories while Schoeni and Blank used four.

γ_s: State fixed effects.

γ_t: Time fixed effects

$trend \bullet \gamma_s$: Linear state-specific time trends

Z: A vector of control variables including current and lagged unemployment rate and current and lagged employment growth rate, each interacted with each of the three education groups; log maximum AFDC/TANF benefit levels for a family of three, interacted with each of the three education groups; education dummies for each of the three groups, year effects interacted with each education group indicator; age dummies for each of the four age groups; each age indicator interacted with each education indicator; proportion Hispanic; and proportion non-Hispanic black.

The identification on policy variables occurs in two ways. First, by controlling for a rich set of variables, including state economic measures, state and year fixed effects, as well as demographic characteristics, the independent effects of the policy variables should be evident. Second, if one believes there are still important omitted variables in this model that bias the coefficients on the policy variables, one can compare the effects of policy among less skilled versus more skilled men (or women) as a test of whether the measured effects make sense. Welfare changes should have few effects on women with more than a high school diploma since so few women with post-high school training receive welfare, and should have few effects on more skilled men as well. Schoeni and Blank indicate that most of their measured effects of policy are present among less skilled women but not among more skilled women, lending credibility to their results.

We calculate labor force measures among males and females based on ORG data within each state-year-age-education cell.[15] We focus on two dependent variables: labor force participation rates and mean hours of work among all workers.

Columns 1 and 2 of Table 3 essentially show that we can replicate the Schoeni and Blank results for women using our ORG-based labor market data. Column 1 shows the impact of the implementation of waiver policies and of TANF on female labor force participation by education group. Column 2 uses hours of work among all female workers as the dependent variable. Waivers, which were largely designed to increase states' ability to run serious welfare-to-work programs, had a positive and significant effect on female labor force participation among women with less than a high school degree. But the effect of these waivers is essentially zero on more educated women. Waivers also have a positive effect on mean hours of work among workers for the least educated.

The implementation of the TANF program also has a positive effect on labor force participation of the least educated women, but the effect is small and not significant. The effect of TANF on hours among the least skilled women is negative (not the expected sign) and marginally significant, certainly suggesting that there were no large hours gains among women due to TANF. (The negative sign could be due to the large increase in the number of less skilled women who entered the labor market after the enactment of TANF, potentially driving down average hours worked.)

[15] Schoeni and Blank use March CPS data, which provides fewer observations and is likely to be noisier data, but they were interested in AFDC receipt, family income and poverty measures as well as labor force participation. These data were available only in the March CPS, which limited them to that data source with fewer observations per cell.

In contrast, columns 3 and 4 show equivalent estimates for men, using labor force participation among men as the dependent variable in column 3 and mean hours of work among male workers in column 4. If these policy changes are primarily affecting labor market decisions among women, then any significant coefficient must reflect the indirect impact of female labor supply changes on male labor force behavior.

In column 3, there is no evidence of any displacement effect of these policy changes on male labor supply. Indeed, the implementation of waivers appears to have had a small but significant positive effect on labor force participation among the least skilled men (and no effects on more skilled men.) At best this suggests there might have been some direct and positive effects of waivers on male labor supply; it certainly provides no evidence of displacement or substitution between male and female workers. There are no effects of the implementation of TANF on male labor force participation among less or more skilled men.

The results on mean hours of work among workers are a bit harder to interpret. Consistent with the effect on male labor force participation, waivers appear to have significantly increased hours of work among men, again not consistent with the substitution hypothesis. The implementation of TANF appears to have reduced male hours of work among the least skilled. It is hard to interpret this as a substitution effect, however, since the least skilled women also show negative hours effects due to TANF. Even more puzzling for the men, the implementation of TANF appears correlated with a strong and positive increase in labor supply among more skilled men. While it may be possible to tell a story justifying such a coefficient (perhaps hiring more low-skilled

women required an increase in more-skilled supervisors), it is clear that these coefficients provide no evidence of displacement or substitution.

In short, the welfare reforms enacted in the 1990s did appear to have some positive effects on female labor supply, especially the waivers. (The effects of TANF on female labor force participation seem relatively minor.[16]) There is no evidence in our estimates that these policies had negative effects on male labor force involvement, however. In fact, the effect of waivers on the labor force participation of the least skilled men is positive, although this is the place where displacement might be expected to be the largest given that this is the policy that has the largest positive effects on labor supply among least skilled women.

5. Changes on the Margin in Labor Force Involvement

The analysis conducted so far focuses on levels or changes in overall rates of labor force participation or hours worked. It seems likely that if there are significant substitution effects, however, they will operate largely along the margin of new hires. That is, they should show up more readily in the flows in and out of employment than they do in overall rates of labor force participation. Average measures of labor force participation or hours like the ones discussed above will respond to changes in the hiring process, so they are useful topics of analysis. However, changes in the flows into and out of employment may be a step higher in the causal chain. These transitions are the topic of this section.

[16] Schoeni and Blank indicate that the economy accounts for much of the female labor force participation increases post 1996. Once economic variables are controlled for, TANF implementation has no further effects.

We utilize March CPS data to examine the rate at which men are transitioning into, say, full-time jobs. The March CPS is particularly useful because it has information both on individual labor force involvement "last week" - that is, in the week prior to the CPS interview - and "last year" - that is, in the previous year. We base our analysis on prime age men, ages 25 to 54, over the years 1989 to 2000. In the regression analysis described below (but not the plots) we also restrict our data to men who live in defined MSAs, since the employment transitions among women within the relevant MSA are a key independent variable.

We begin by considering the probability that a man is in full-time employment last week, conditional on whether he was or was not typically employed full-time the previous year on his usual job.[17] Full-time employment is defined as working at least 35 hours last week (or working at least 35 hours per week last year in one's usual job). Figure 4 plots the fraction of men who were employed full-time last week conditional on their employment status last year. The figure shows that full-time prime-age men faced layoffs during the recession of the early 1990s. By contrast, the probability of working full-time last week among men who did not usually work full-time last year is much lower and does not appear very sensitive to the business cycle.

Figures 5 and 6 show similar plots among two subpopulations of particular interest, male high school dropouts and Black men. These figures exhibit the same basic pattern as the plot for all men, though the probability of full-time work in the past week is lower for both of these subgroups.

[17] Men who do not work full-time include both those working part-time on their usual weekly job and those who are out of the labor market all of the last year. Relatively few men do not report any hours per week on at least one job in the previous year; 16 percent of high school dropouts, 7.5 percent of high school graduates, and 4 percent of those with more than a high school degree have no reported work last year.

A second outcome of interest is the probability that a man worked at all last week, conditional on his attachment to the labor force last year. We consider men who worked fewer than 13 weeks last year relatively unattached to the labor force, while men who worked at least 13 weeks are considered attached to the labor force. Figures 7, 8, and 9 plot rates of transition into employment last week conditional on last year's labor force attachment level among all men in the sample (Figure 7), high school dropouts only (Figure 8) and Black men only (Figure 9).

These figures are similar to those for the full-time/not-full-time outcomes. They indicate that more attached men are affected by the business cycle of the early 1990s. Men who were less attached to the labor force in the previous year show fewer cyclical effects but are much less likely to be employed at any level in the previous week.

If increased female employment is crowding out increased male employment via increased hiring of women relative to men then one would expect male rates of transition into employment to be smaller when female transition rates are greater. Of course, simple time-series evidence is of limited use here given that men and women are both affected by economic conditions. We therefore again use cross-sectional and over-time variation at the MSA level to investigate the relationship between male and female transition rates. Our basic approach will be to create separate samples for all men reporting a particular employment status last year, i.e., a sample of men who worked less than full-time last year and a sample of men who worked full-time last year. For each man, we define a dummy variable indicating the probability of full-time employment status last week. Within each sample we regress this conditional employment probability on a set of explanatory variables. The main variable of interest will be the

average transition rate for women within an MSA. This is the average probability within the MSA that women are now working full-time, conditional upon the same starting labor market status as for the men. Hence, the regression on the sample of men who were not working full-time in the last year uses as an independent variable "the probability within the same MSA that women who were not working full-time last year now work full-time." Similarly, the regression for the sample of men who did work full-time last year uses as a dependent variable "the probability within the same MSA that women who worked full-time last year continue to work full-time."[18]

Controlling appropriately for local changes in economic conditions is an important but difficult task. If economic differences across MSAs in underlying transition rates were fixed over time, we could simply use MSA fixed effects. Nationwide economic changes that affect all employment transitions in a given year could be absorbed using year dummies. But, as before, it is reasonable to believe that the impact of changes in female transition rates on male transitions differs across labor markets. In particular, area-level differences in male and female occupational structure could confound any crowding-out effects of female transitions. MSA-level fixed effects will not solve this problem if the impact of changes in female transitions is greater when, say, gender differences in industrial structure are smaller.

We address this concern by interacting female transition rates with the MSA-level industry segregation index used in Section 3. With segregation of male and female workers across industries, we would expect male transition rates to be less affected by

[18] Obviously, men's transition rates between any one employment state and any other will be associated with the other transition rates. Hence the specifications estimated here are reduced form in nature, so that we have already solved out the full model. A more complete econometric discussion is available on request to the authors.

female transition rates in highly segregated areas. Since the expected effect of female transitions under the substitution hypothesis is to reduce transitions into employment for men, the expected sign of the coefficient on the interaction term is positive.

Other variables included in these regressions are men's race (Black or White) and Hispanic background, their age, and their educational attainment, as well as average MSA-level values of these variables for women. The gender segregation index is also interacted with all of these women's average characteristics.[19] Also included are the one-year-lagged values of the unemployment rate and its interaction with the segregation index. Lastly, a set of year dummies and their interactions with the segregation index are included. In order to obtain reasonably precise estimates of MSA-level average variables for women, we create these averages across all women in the sample at all education levels.

We use data on the same 208 MSAs used in Section 3 and utilize all of the data on men and women within these MSAs. Our data are based on the years 1989 to 2000 (hence, data for "last year" are for calendar years 1988 to 1999). The resulting sample has 234,308 observations on all men.

The results based on the probability of working full-time last week conditional on the hours per week worked in the last year are presented by educational attainment in Table 4 and by race/ethnicity in Table 5. Among dropouts and high school graduates, the point estimates in Table 4 on the "Probability of full-time work among women" variable are negative for only one of the four transitions into full-time employment among less skilled workers. This suggests that, in a completely unsegregated MSA, more frequent

transitions into full-time status for women are associated with more frequent transitions for men--exactly the opposite of the substitution hypothesis. The sign on the interaction with the segregation index is positive in two cases and negative in the other two, suggesting no systematic impact in more or less segregated MSAs. Moreover, the coefficient estimates are generally small and are all statistically insignificant.[20] For the group having more than a high school diploma, the coefficient estimate for the "main effect" are again of mixed sign, while the interaction effects are positive but statistically insignificant.

The results in Table 5, stratified by race and ethnicity, are basically similar. Here the coefficient on the "Probability of full-time work among women" variable is negative in three of the four specifications for less skilled workers, but again none of these estimates is statistically significant. The coefficient estimates for the interaction term are generally large, but they are never statistically significant.

Tables 6 and 7 show the probability of working at all among men, conditional on their weeks of labor market involvement in the last year.[21] Table 6 estimates these transitional probabilities by skill group and Table 7 estimates them by race and ethnicity. In six of the twelve specifications, these two tables show significant effects of "Probability of any employment last week among women". However, in all six cases, the coefficient estimate is positive -- again suggesting the opposite of the substitution hypothesis. The interaction effects are also significant in several cases, but only when

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[19] In econometric terms, one can write down analogous gender-specific reduced form models for transition rates. The underlying algebra shows that one must also control for interactions between average women's characteristics and the segregation index.

[20] Standard errors were estimated using a robust method to account for correlation of unobservables over time within MSAs.

they are negative, again the opposite of the substitution hypothesis. The only cases in which the coefficient estimates are consistent with the substitution hypothesis are for dropouts, Blacks, and Hispanics (conditional on having low workforce attachment last year). But in all cases, these estimates are below standard levels of statistical significance.[22]

All in all, the results in this section suggest no evidence that women crowd out men in transitions to full-time employment (conditional on last year's full-time status), or to employment (conditional on last year's labor force attachment.) While it can be very difficult to infer causality from regressions like these, the estimates suggest that any crowdout effects are dominated by trends in local labor markets.

6. Conclusions

This paper has searched for evidence of substitution over the 1990s between male and female workers. All of the estimation techniques that we use here have some interpretational difficulties; in no case do we want to claim that we have clearly estimated causal effects. Yet we cannot find any evidence of displacement or substitution between less skilled male and female workers in the labor market.

To a larges extent, the discussion in this chapter underscores the difficulties inherent in studying the substitution hypothesis empirically. On one hand, one would have expected that crowd-out effects, should they exist, would have been discernible during the last business cycle: Both the strong secular trend in female labor force

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[21] The dependent variable in Tables 6 and 7 (and the key independent variable on women as well) is now defined as “employed last week”, and not full-time employment.

[22] We estimated the same models, instead defining strong labor force attachment by having worked at least 26 weeks last year. The results are qualitatively identical to those reported in Tables 6 and 7.

participation and enormous reductions in the welfare rolls should have provided enough increases in female competitors to observe crowd-out. On the other hand, macroeconomic performance in the last several years of the 1990s -- when the welfare rolls were plummeting -- was unprecedented in its strength. Untangling these two effects is extremely difficult.

Even with these problems, however, these estimates suggest than any substitution that occurred was probably relatively minor. The techniques we use here are likely to have uncovered evidence of female/male substitution if those effects were substantial in metropolitan labor markets. While we cannot entirely rule out the possibility for displacement and substitution among less skilled male and female workers with this analysis, we can say that any effect that did occur was relatively small and clearly not sufficient to provide an adequate explanation for the declines in male labor force participation observed among less skilled male workers in the 1990s.

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Table 1

Simple Correlations Between Male and Female Labor Supply

	Education Level			
	All	**Less Than High School**	**Exactly High School**	**Low Skill (High School or Less)**
Correlations:				
1. Change in male LFP rates versus change in female LFP rates within skill level, 1989-2000.	34.4	26.2	30.4	26.2
2. Change in mean hours among male workers versus change in female LFP rates within skill level, 1989-2000.	-4.9	-5.7	-4.8	-14.3
Difference-in-Difference Calculations:				
3. Change in male LFP rates from 1989-2000 in 20 MSAs with highest rate of LFP growth among less-skilled women.	-0.004	0.042	-0.018	0.019
4. Change in male LFP rates from 1989-2000 in 20 MSAs with lowest rate of LFP growth among less-skilled women.	-0.042	-0.072	-0.053	-0.055
5. Difference in differences	0.038	0.114	0.035	0.074

Table 2
Effects of Changes in Female Labor Force Participation on Male Labor Supply
1989 to 2000 in 208 MSAs

Part I: Dependent Variable = Change in Labor Force Participation Rate Among Men

Included Variables:	All Men (1)	(2)	(3)	Less Skilled (HS or less) (4)	(5)	(6)	Less than HS Only (7)	(8)	(9)
Change in Female LFP[a]	.218 (.020)	.215 (.020)	.220 (.020)	.169 (.023)	.163 (.023)	.164 (.022)	.191 (.041)	.175 (.041)	.177 (.041)
MSA Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Unemployment Rates[b]	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes
Other Control Variables[c]	No	No	Yes	No	No	Yes	No	No	Yes

Part II: Dependent Variable = Change in Mean Hours of Work Among Male Workers

Included Variables:	All Men (1)	(2)	(3)	Less Skilled (HS or less) (4)	(5)	(6)	Less than HS Only (7)	(8)	(9)
Change in Female LFP[a]	-.537 (.728)	-.583 (.730)	-.565 (.731)	.497 (.766)	.519 (.769)	.540 (.770)	-.516 (1.963)	-.566 (1.970)	-.517 (1.975)
MSA Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Unemployment Rates[b]	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes
Other Control Variables[c]	No	No	Yes	No	No	Yes	No	No	Yes

Standard errors in parentheses.
Number of observations = 2496 (208 MSAs for 12 years)

[a] Change in female labor force participation (LFP among all women used in columns (1) through (3). Change in female LFP among all low skilled women (women with less than a HS degree) used in columns (4) through (9).
[b] Includes current and lagged unemployment rates by MSA
[c] Includes MSA-specific variables for each year for percent black, percent Hispanic, percent employed in manufacturing, percent employed in nonprofessional services, percent employed in professional services; percent employed in public administration, and gender industry segregation index.

Table 3
Effects of Welfare Waivers and TANF on Male and Female Labor Supply Variables,
1979-2000

Policy Effects (Interacted with Skill)	Women		Men	
	Labor Force Participation	Mean Hours (Among Wkrs)	Labor Force Participation	Mean Hours (Among Wkrs)
Waiver • Educ<12	.024	.273	.016	.302
	(.006)	(.172)	(.004)	(.172)
Waiver • Educ=12	.002	.179	-.001	-.064
	(.004)	(.129)	(.003)	(.137)
Waiver • Educ>12	-.004	-.094	-.002	.242
	(.004)	(.108)	(.003)	(.111)
TANF • Educ<12	.011	-.649	.001	-.864
	(.014)	(.421)	(.011)	(.422)
TANF • Educ=12	-.001	.192	.004	.038
	(.011)	(.337)	(.009)	(.356)
TANF • Educ>12	.006	.104	.002	.618
	(.009)	(.263)	(.007)	(.272)

Standard errors in parentheses. Regressions weighted by population in each cell. Observations are defined by averages within each state-year-education-age group for 51 states across the 12 years indicated.

Each column represents a regression including current and lagged unemployment rate and employment growth rate interacted with each education group, maximum benefit for a family of three interacted with education groups, age dummies (16-25, 26-44, 45-64), education dummies, age dummies interacted with education dummies, race (Hispanic, non-Hispanic black), year effects, year effects interacted with education dummies, state effects, and state-specific time trends.

Table 4
Estimated Probability of Full-time Employment Among Men in March, Conditional on Employment Status Last Year: By Education Level

	Job status last year	
	Not Full-time	Full-time
Education: Less than HS		
Mean of LHS variable	0.11	0.76
Probability of Full-time Work Among Women[a]	0.07 (0.67)	0.46 (0.35)
Ind Segregation Index • Prob of Full-time Work Among Women	0.14 (1.96)	-0.82 (1.01)
Number of Observations	6,916	23,283
Education: HS Only		
Mean of LHS Variable	0.13	0.82
Probability of Full-time Work Among Women[a]	-0.61 (0.61)	0.11 (0.21)
Ind Segregation Index • Prob of Full-time Work Among Women	2.14 (1.76)	-0.01 (0.62)
Number of Observations	8,070	58,210
Education: More than HS		
Mean of LHS Variable	0.16	0.86
Probability of Full-time Work Among Women[a]	-0.14 (0.66)	0.06 (0.14)
Ind Segregation Index • Prob of Full-time Work Among Women	1.23 (2.01)	0.24 (0.42)
Number of Observations	10,491	109,125

Standard errors in parentheses are robust to arbitrary-form within MSA-year cells.
* Indicates significance at the 0.10 level; ** at the 0.05 level; *** at the 0.001 level.
All specifications include demographic covariates and year fixed effects.
[a]Average probability of full time work last week among all women within an MSA, conditional on zero or part-time work last year (Column 1 estimates) or full-time work last year (column 2 estimates). This average includes all women in all educational groups.

Table 5
Estimated Probability of Full-time Employment Among Men in March, Conditional on Employment Status Last Year: By Race and Ethnicity

	Job status last year	
	Not Full-time	Full-time
Black Men		
Mean of LHS variable	0.10	0.79
Probability of Full-time Work Among Women[a]	-0.81 (0.90)	-0.36 (0.36)
Ind Segregation Index • Prob of Full-time Work Among Women	3.12 (2.65)	1.46 (1.05)
Number of Observations	4,499	16,948
Non-Hispanic White Men		
Mean of LHS Variable	0.14	0.84
Probability of Full-time Work Among Women[a]	-0.34 (0.43)	0.14 (0.14)
Ind Segregation Index • Prob of Full-time Work Among Women	1.44 (1.26)	0.01 (0.42)
Number of Observations	14,296	134,364
Hispanic Men		
Mean of LHS Variable	0.16	0.81
Probability of Full-time Work Among Women[a]	0.77 (1.06)	0.53 (0.33)
Ind Segregation Index • Prob of Full-time Work Among Women	-1.78 (3.27)	-1.19 (0.96)
Number of Observations	5,229	30,958

Standard errors in parentheses are robust to arbitrary-form within MSA-year cells.
* Indicates significance at the 0.10 level; ** at the 0.05 level; *** at the 0.001 level.
All specifications include demographic covariates and year fixed effects.
[a]Average probability of full time work last week among all women within an MSA, conditional on zero or part-time work last year (Column 1 estimates) or full-time work last year (column 2 estimates). This average includes all women in all race/ethnic groups.

Table 6
Estimated Probability of Any Employment Among Men in March, Conditional on Number of Weeks Worked Last Year: By Education Level

	Weeks worked last year	
	No more than 13	At least 13
Education: Less than HS		
Mean of LHS variable	0.16	0.89
Probability of Employment Among Women[a]	-0.49	0.93**
	(0.38)	(0.41)
Ind Segregation Index • Prob of Employment Among Women	1.31	-2.19*
	(1.11)	(1.17)
Number of Observations	5,787	24,412
Education: HS Only		
Mean of LHS Variable	0.18	0.92
Probability of Employment Among Women[a]	0.38	0.40*
	(0.50)	(0.22)
Ind Segregation Index • Prob of Employment Among Women	-1.30	-0.90
	(1.49)	(0.66)
Number of Observations	6,391	59,889
Education: More than HS		
Mean of LHS Variable	0.22	0.95
Probability of Employment Among Women[a]	1.00*	0.11
	(0.57)	(0.12)
Ind Segregation Index • Prob of Employment Among Women	-2.96*	-0.25
	(1.69)	(0.36)
Number of Observations	6,539	113,077

Standard errors in parentheses are robust to arbitrary-form within MSA-year cells.
* Indicates significance at the 0.10 level; ** at 0.05 level; *** at the 0001 level.
All specifications include demographic covariates and year fixed effects.
[a]Average probability of any employment last week among all women within an MSA, conditional on less than 13 weeks last year (Column 1 estimates) or at least 13 weeks last year (column 2 estimates). This average includes all women in all educational groups.

Table 7
Estimated Probability of Any Employment Among Men in March, Conditional on Number of Weeks Worked Last Year: By Race and Ethnicity

	Weeks worked last year	
	No more than 13	More than 13
Black Men		
Mean of LHS variable	0.15	0.89
Probability of Employment Among Women[a]	-0.04 (0.66)	0.98* (0.54)
Ind Segregation Index • Prob of Employment Among Women	0.62 (2.03)	-2.38 (1.63)
Number of Observations	3,862	17,585
Non-Hispanic White Men		
Mean of LHS Variable	0.18	0.94
Probability of Employment Among Women[a]	0.72** (0.34)	0.21* (0.12)
Ind Segregation Index • Prob of Employment Among Women	-2.40** (1.02)	-0.53 (0.36)
Number of Observations	9,894	138,766
Hispanic Men		
Mean of LHS Variable	0.22	0.92
Probability of Employment Among Women[a]	-0.31 (0.96)	0.36 (0.32)
Ind Segregation Index • Prob of Employment Among Women	0.96 (3.05)	-0.64 (0.93)
Number of Observations	3,915	32,272

Standard errors in parentheses are robust to arbitrary-form within MSA-year cells.
* Indicates significance at the 0.10 level; ** at 0.05 level; *** at the 0001 level.
All specifications include demographic covariates and year fixed effects.
[a]Average probability of any employment last week among all women within an MSA, conditional on less than 13 weeks last year (Column 1 estimates) or at least 13 weeks last year (column 2 estimates). This average includes all women in all race/ethnic groups.

Figure 1

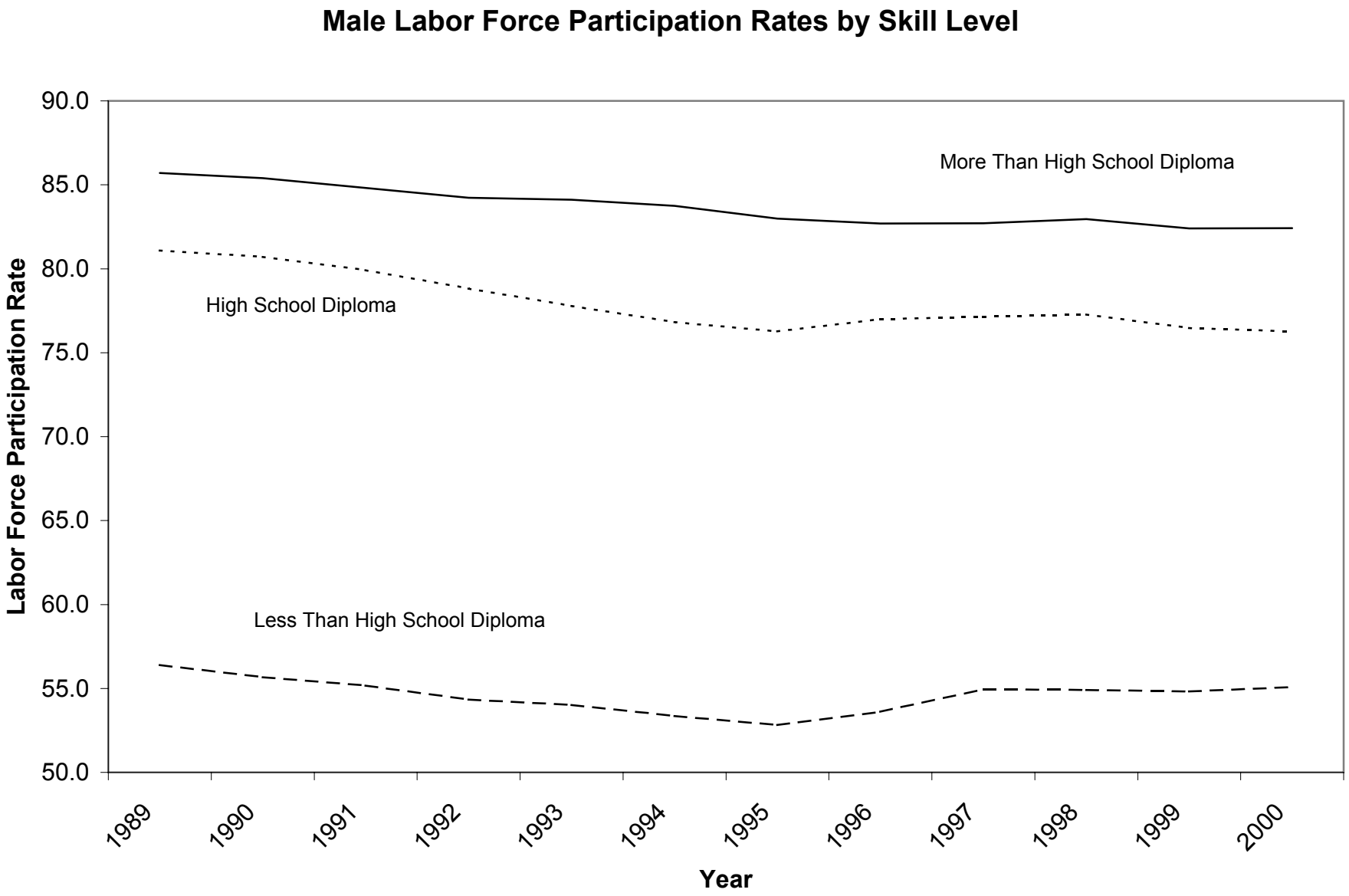
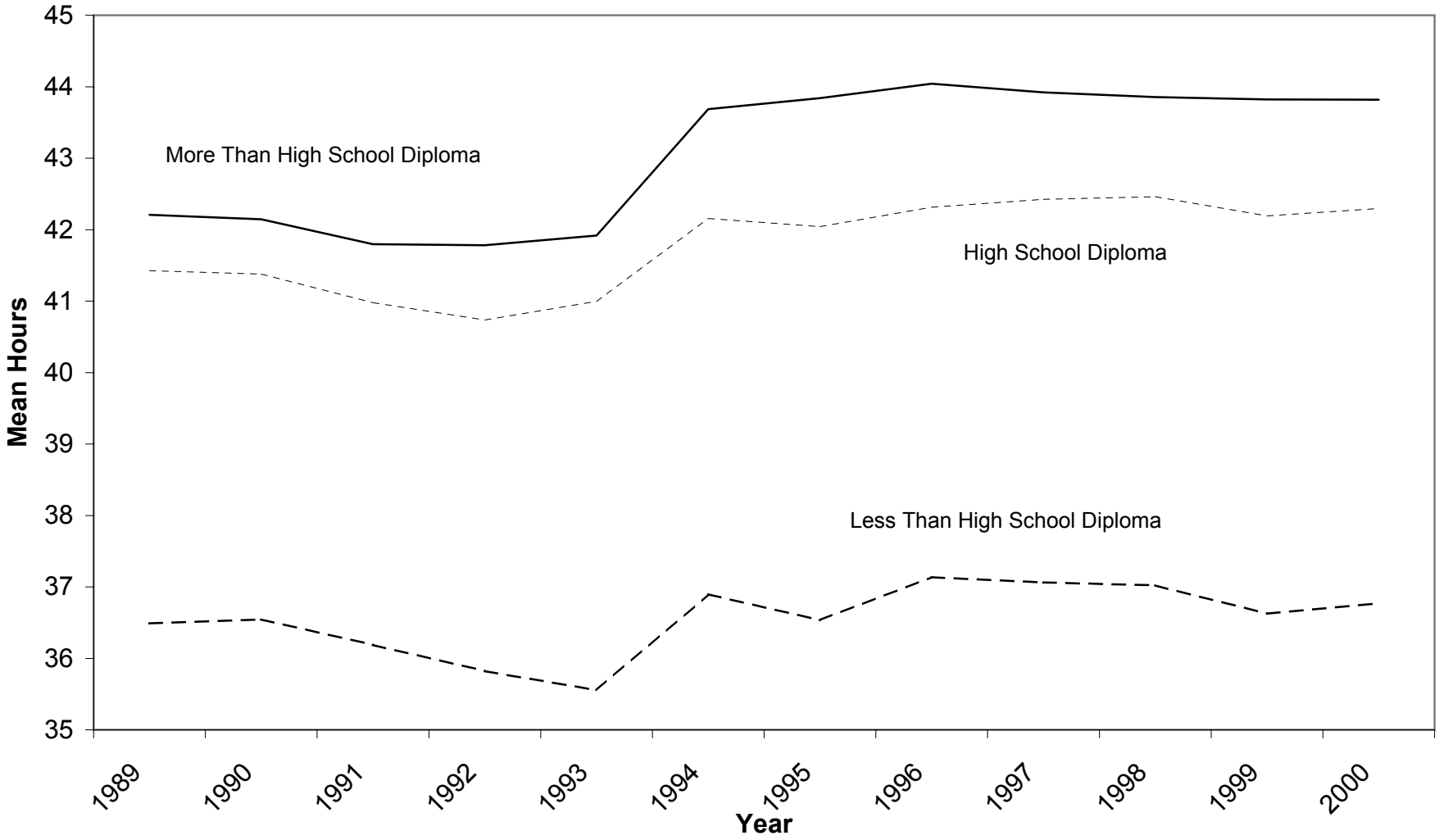

Male Labor Force Participation Rates by Skill Level
90.0
85.0
80.0
75.0
70.0
65.0
60.0
55.0
50.0
Labor Force Participation Rate
More Than High School Diploma
High School Diploma
Less Than High School Diploma
1989
1990
1991
1992
1993
1994
1995
1996
1997
1998
1999
2000
Year

Figure 2

Mean Hours of Work Among Men Who Work by Skill Level



Note: Definitional break in the data between 1993 and 1994.

Figure 3

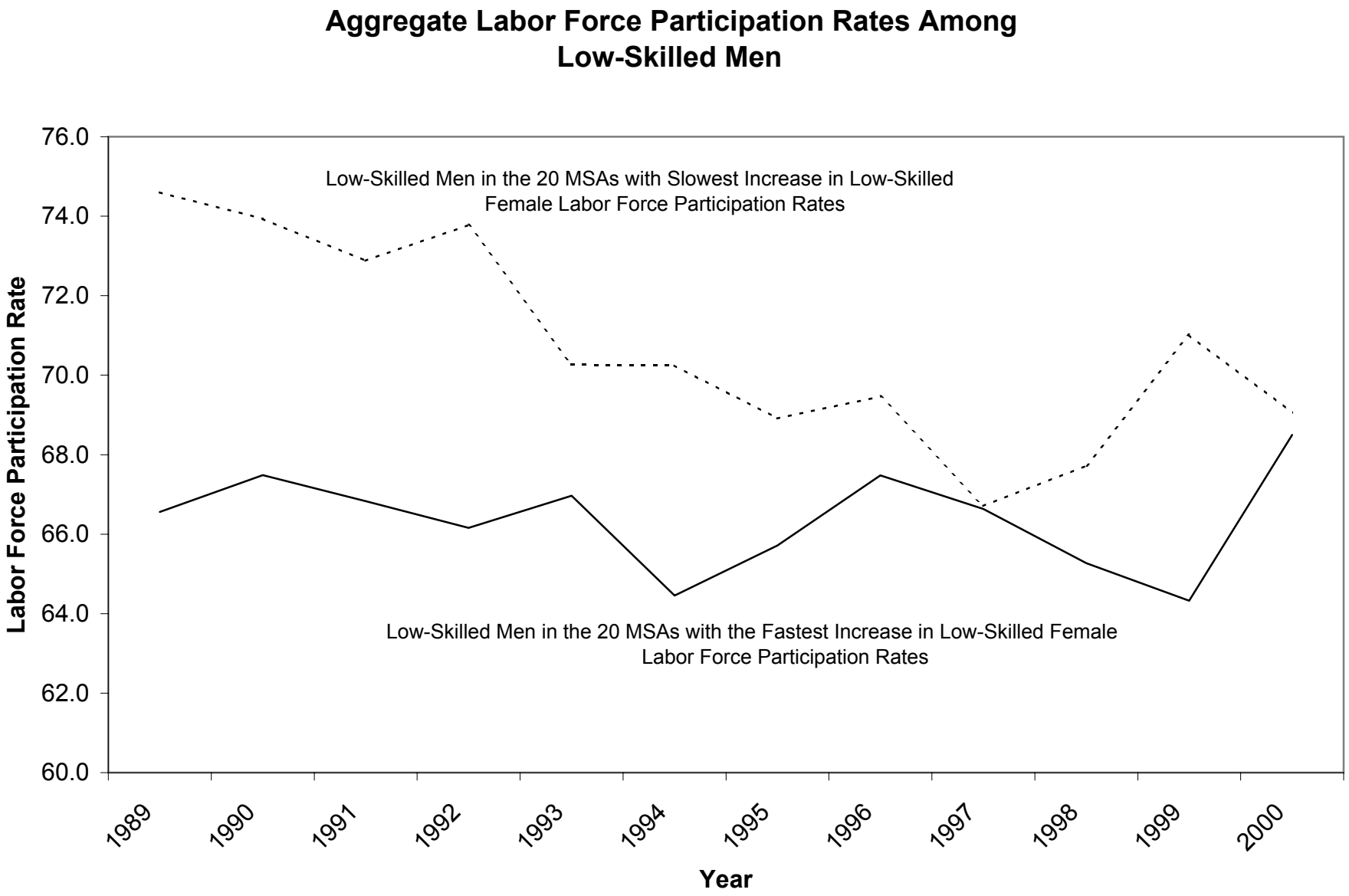

Aggregate Labor Force Participation Rates Among Low-Skilled Men
Low-Skilled Men in the 20 MSAs with Slowest Increase in Low-Skilled Female Labor Force Participation Rates
Low-Skilled Men in the 20 MSAs with the Fastest Increase in Low-Skilled Female Labor Force Participation Rates
Labor Force Participation Rate
76.0
74.0
72.0
70.0
68.0
66.0
64.0
62.0
60.0
1989
1990
1991
1992
1993
1994
1995
1996
1997
1998
1999
2000
Year

Figure 4

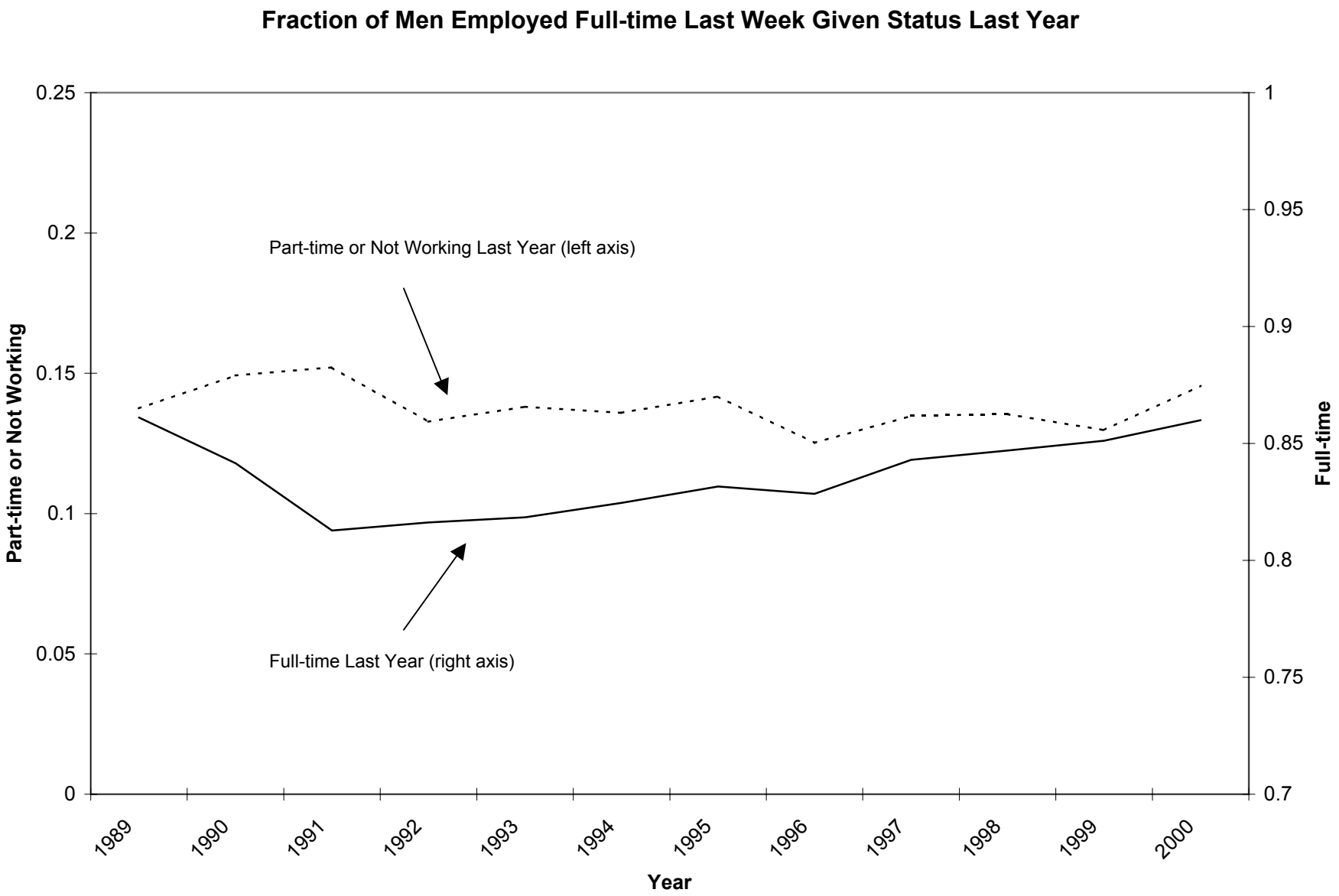

Fraction of Men Employed Full-time Last Week Given Status Last Year
0.25
0.2
0.15
0.1
0.05
0
1
0.95
0.9
0.85
0.8
0.75
0.7
Part-time or Not Working
Full-time
Part-time or Not Working Last Year (left axis)
Full-time Last Year (right axis)
1989
1990
1991
1992
1993
1994
1995
1996
1997
1998
1999
2000
Year

Figure 5

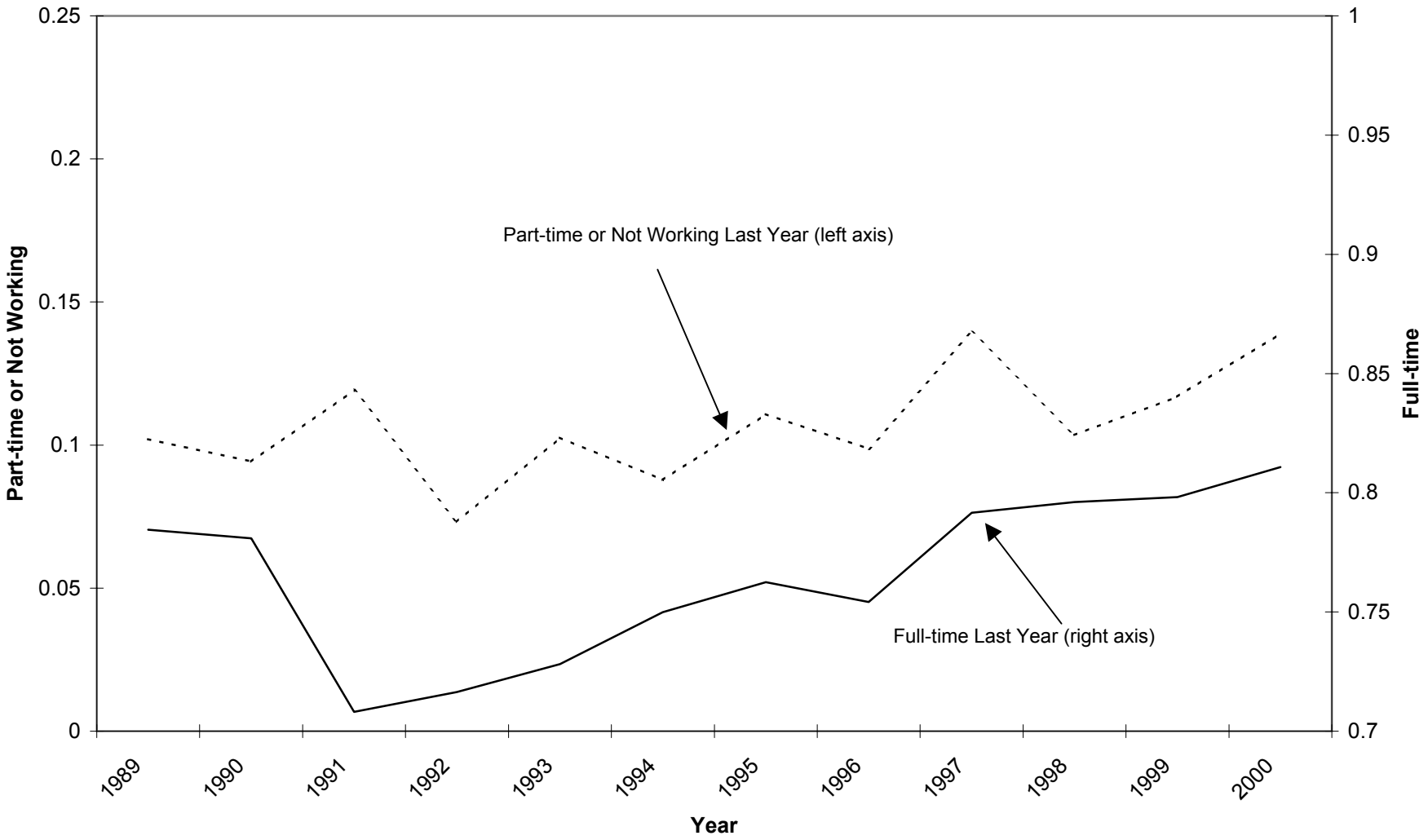

Fraction of Men Emplyed Full-time Last Week Given Status Last Year,
Less Than High School Degree Only
Part-time or Not Working
Full-time
Year
Part-time or Not Working Last Year (left axis)
Full-time Last Year (right axis)
0.25
0.2
0.15
0.1
0.05
0
1
0.95
0.9
0.85
0.8
0.75
0.7
1989
1990
1991
1992
1993
1994
1995
1996
1997
1998
1999
2000

Figure 6

Fraction of Men Employed Full-time Last Week Given Status Last Year, Black Men Only

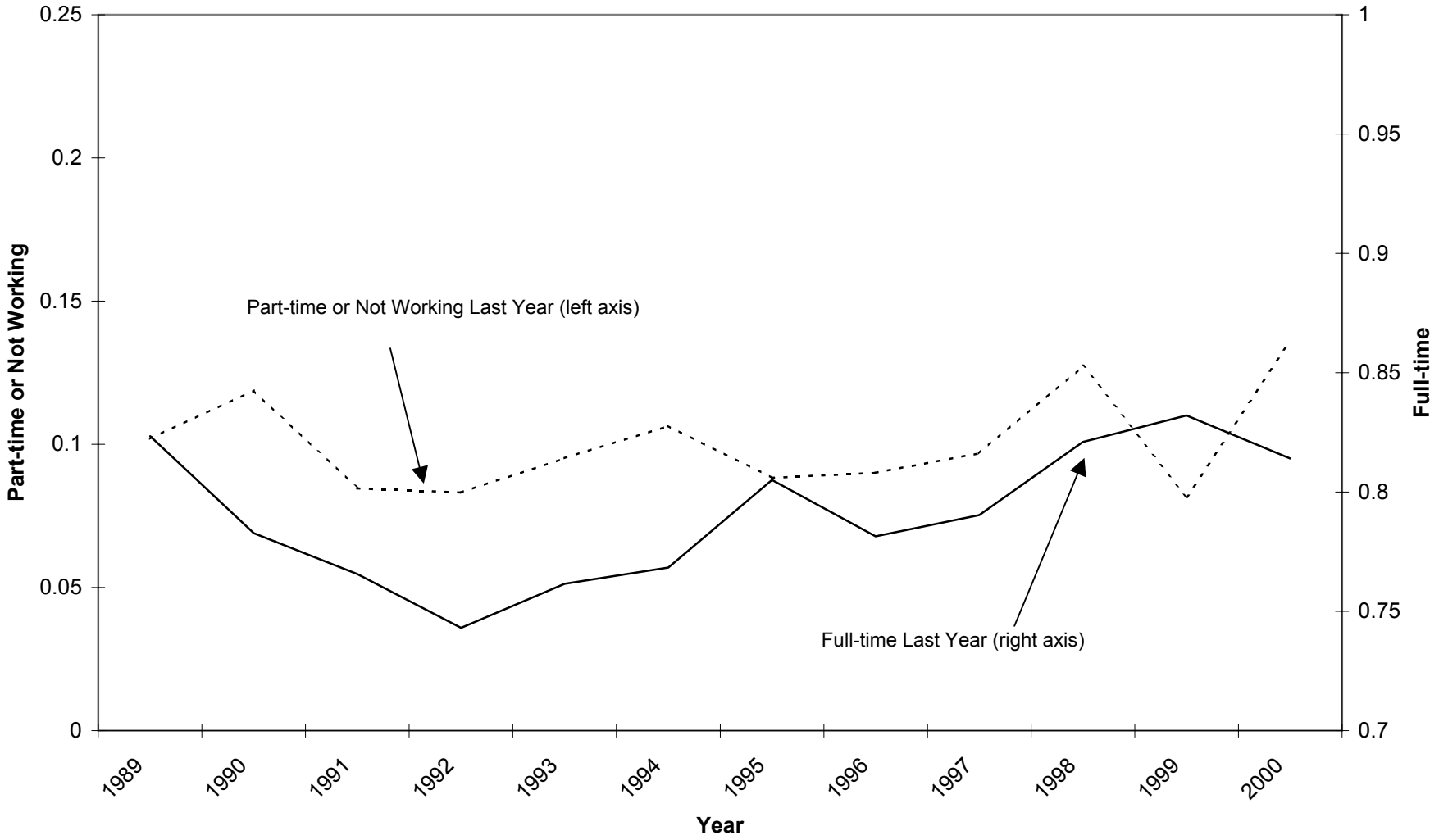


Figure 7

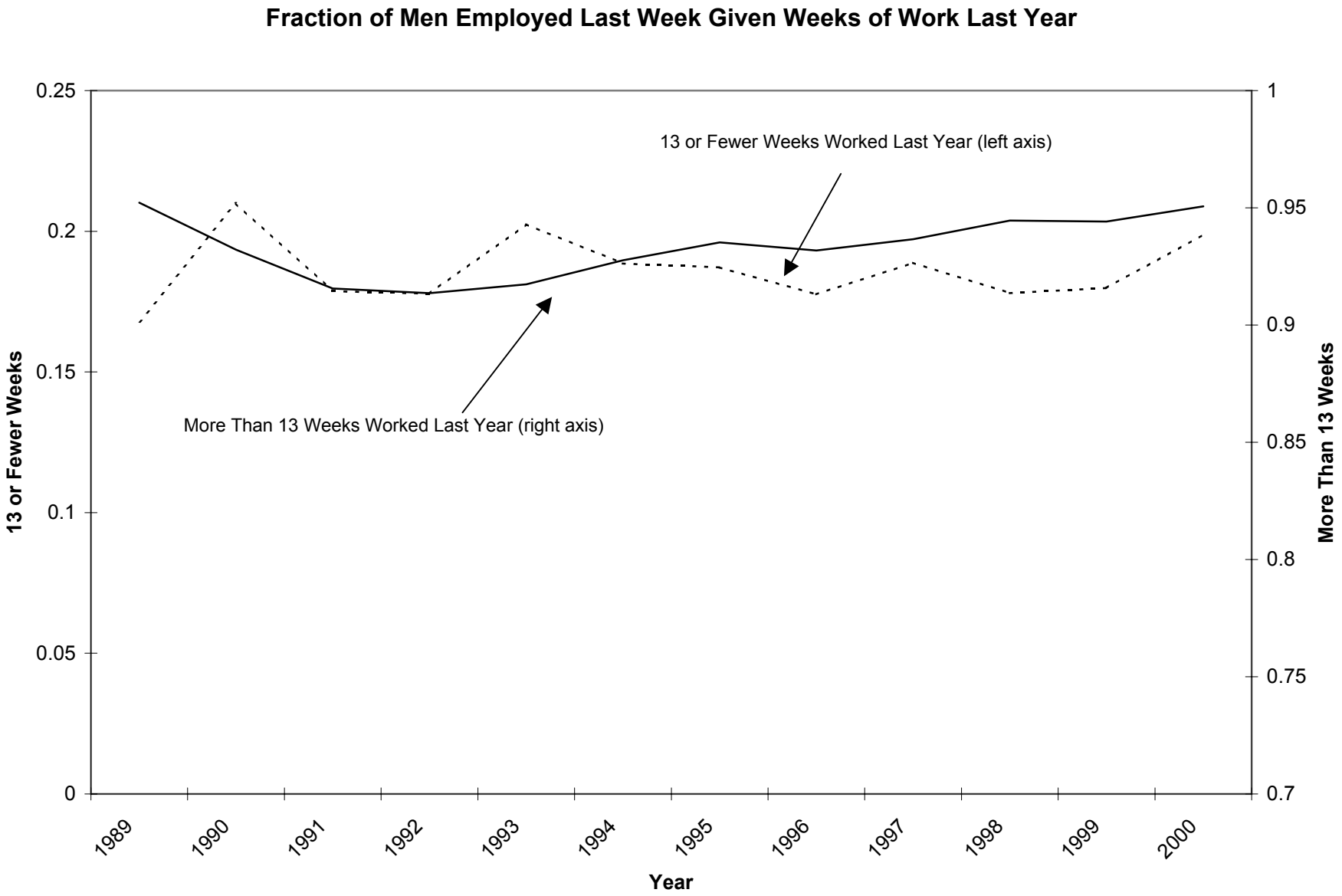

Fraction of Men Employed Last Week Given Weeks of Work Last Year
13 or Fewer Weeks Worked Last Year (left axis)
More Than 13 Weeks Worked Last Year (right axis)
13 or Fewer Weeks
More Than 13 Weeks
Year
0.25
0.2
0.15
0.1
0.05
0
1
0.95
0.9
0.85
0.8
0.75
0.7
1989
1990
1991
1992
1993
1994
1995
1996
1997
1998
1999
2000

FIgure 8

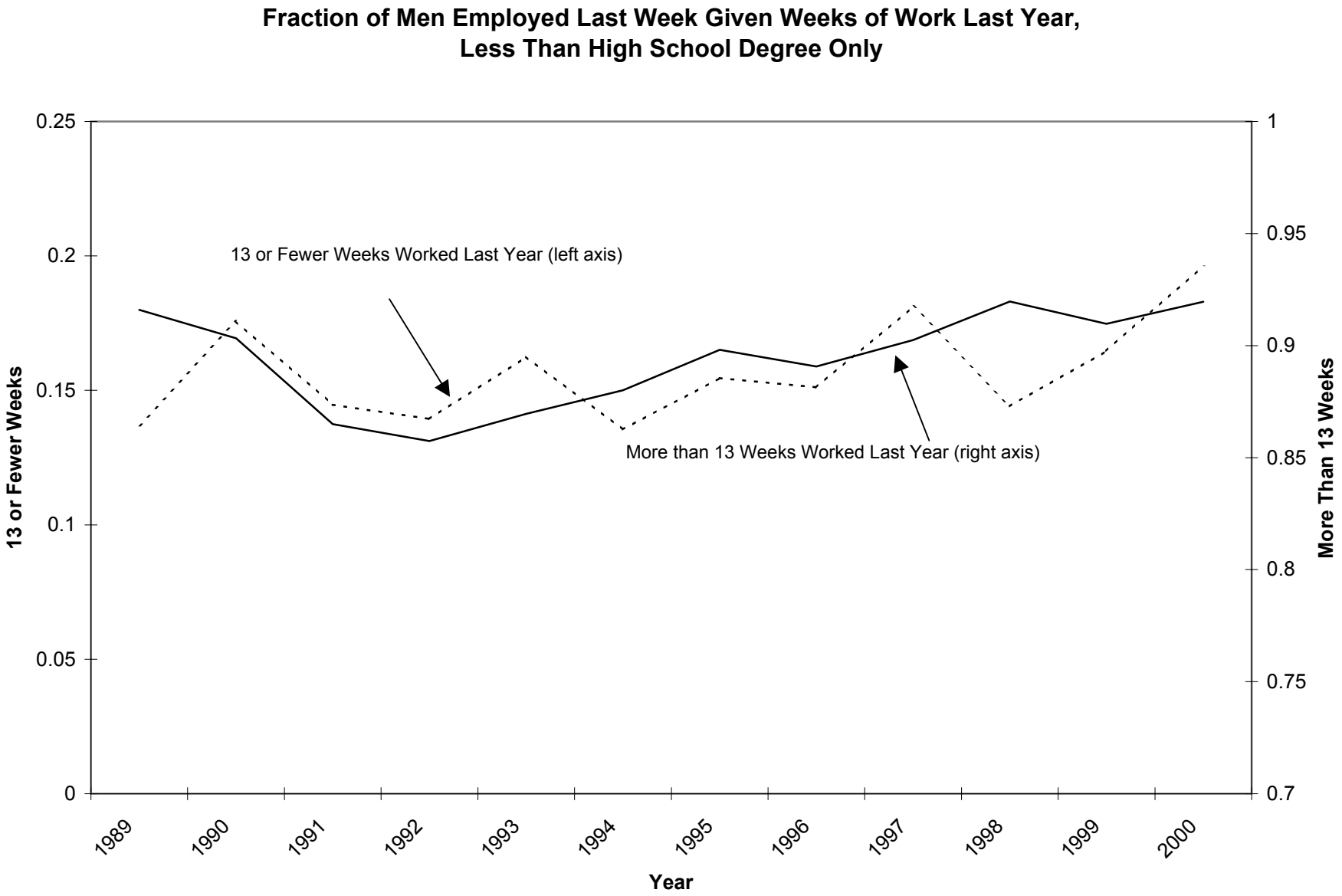

Fraction of Men Employed Last Week Given Weeks of Work Last Year,
Less Than High School Degree Only
13 or Fewer Weeks Worked Last Year (left axis)
More than 13 Weeks Worked Last Year (right axis)
13 or Fewer Weeks
More Than 13 Weeks
Year
0.25
0.2
0.15
0.1
0.05
0
1
0.95
0.9
0.85
0.8
0.75
0.7
1989
1990
1991
1992
1993
1994
1995
1996
1997
1998
1999
2000

Figure 9

Fraction of Men Employed Last Week Given Weeks of Work Last Year, Black Men Only

