Monetary Tightening and the Dynamics of US Race and Gender Stratification

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Abstract. This article explores the race and gender effects of monetary tightening in the US from 1979–2008 using state-level panel data. Results indicate the costs of fighting inflation are unevenly distributed amongst workers, weighing more heavily on black females and black males, followed by white females, and lastly white males. We also find evidence that the relative unemployment costs of monetary tightening for subordinate groups vary with the black share of the population.

Introduction

Central banks across the globe have shifted the emphasis of monetary policy to an almost singular concern with controlling inflation over the goal of employment generation.[1] The primary instrument in the central banker's toolkit is nominal interest rates, designed to act on the demand side of the economy by slowing consumption and investment. A potential cost of controlling inflation via this method is an increase in unemployment.

Inflation targeting has been criticized on several grounds. In developing countries especially, inflation tends to be the result of supply-side bottlenecks rather than excess aggregate demand. That structural feature of developing countries inhibits the effectiveness of inflation

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targeting and raises the social costs of reducing inflation. Further, even when inflation is demand-induced, a body of evidence suggests that inflation rates below 15–20 percent are not harmful to growth, suggesting central banks could do much more to reduce unemployment than they currently are doing (Pollin and Zhu 2006).

A third concern, one we explore here, is that the costs of fighting inflation are unevenly distributed. Disinflationary monetary policy can exacerbate gender and racial inequalities if subordinate groups experience a disproportionate share of the resulting job losses, which induce competition over scarce jobs. In this article, we explore whether black men, black women, and white women fare worse in the competition over jobs relative to white men in response to disinflationary policy. Unequal effects may be transmitted indirectly in ways that reflect structural features of gender and racial hierarchies. In the US context, men and women of color and white women tend to be concentrated in precarious forms of employment that face a greater likelihood of elimination when demand contracts. Racial and gender effects of disinflationary policy also result from social stratification, whereby norms and stereotypes identify men, and in particular, white men as more deserving of jobs when jobs are scarce, ratifying both gender and racial hierarchies.

Prior empirical investigations have separately investigated either race or gender effects of contractionary monetary policy. A consistent finding is that African Americans bear a heavier burden of joblessness relative to whites in response to interest rate hikes. In contrast, the gender-focused research has yielded contradictory results on whether women's unemployment responds disproportionately to monetary tightening.

Because published work has not considered the interaction of race and gender hierarchies, little is known about the combined effects of race and gender discrimination, or about job competition between subordinate groups in response to monetary policy-induced economic slowdowns. Our study contributes to the literature by explicitly assessing the extent of multiple discrimination, utilizing white males as our reference group.[2]

The goals of the empirical analysis are three-fold. First, we seek to determine whether the methodological approach we utilize yields results consistent with previous research indicating that the costs of contractionary monetary policy are unevenly distributed between ethnic groups. Our approach differs in that white males represent the dominant group, while previous studies include all whites as the reference group. Second, we capitalize on the variation of the share of African Americans in the population at the state level to explore the possibility that in cases of monetary tightening, race is a more salient factor in allocating scarce jobs than gender. We do this by evaluating the effect of black population density on the female/white male unemployment ratio, hypothesizing that the size of the effect is inversely related to black population share.

Third, we explore the data to determine whether, at critical levels, black population density triggers a shift in white attitudes consistent with either threat or contact theory. Threat theory postulates that increases in black population density can intensify racialist group identity in response to whites' perceived threat to their group position. This could generate intensification of racial norms and stereotypes, resulting in blacks bearing a disproportionate burden of the increase in unemployment resulting from contractionary monetary policy. Conversely, contact theory suggests that greater contact, measured as black population density, weakens the propensity for discrimination on the part of whites, resulting in a lower unemployment rate gap between blacks and whites.

Anticipating the results of this analysis, we find that white women, black women, and black men are more likely to experience increases in unemployment than white men in response to disinflationary monetary policy. Those effects are more negative for black women and men than white women. We also find evidence that the relationships between relative unemployment rates and our interest rate policy variable vary with the black share of the population. These findings are important for macroeconomic policymaking and, more specifically, monetary policy. The evidence underscores that macroeconomic policy is neither race- nor gender-neutral. Apart from the inherent problem of socially distortionary policies, recent scholarship shows that lack of attention to distributional effects of macroeconomic policies can produce negative long-run consequences for the economy in terms of lost productivity.

The Distributional Effects of Contractionary Monetary Policy

Recent decades have witnessed a shift in central bank policy from a dual concern with both employment and inflation to an almost exclusive focus on keeping inflation low and close to zero (Epstein and Yeldan 2008). The change in policy emphasis has occurred in both developed and developing economies. The distributional effects of inflation targeting are of great interest in the context of widening income and wealth gaps within and between countries over the last three decades (ILO 2008). Here we focus on employment outcomes as one of the central ways in which monetary policy impacts inequalities in income and economic opportunity.

The primary tool used in inflation targeting is the manipulation of short-term interest rates (in the US, the federal funds rate) charged to banks. Interest rate changes are intended to work on the demand side of the economy. In the US, an increase in the real federal funds rate raises the cost of lending to banks, thereby reducing borrowing for investment and consumption. The effects of contractionary policy on employment are summarized in the concept of the sacrifice ratio, measured as the percentage decline in employment (alternatively, output) in response to a one percent decline in the rate of inflation.

A critical question is whether the impacts of interest rate-induced economic contractions vary systematically by gender and race. William Grieder (1987), in a series of interviews with former Federal Reserve Bank members of the Board of Governors, found they believed their policies to be distributionally neutral and their decisions, rather than rewarding one group or another, simply pursued their vision of sound macroeconomic management. Abell (1991) argues that although Federal Reserve reaction functions appear to only emphasize aggregate concerns—price stability, unemployment rates, and interest rates—the sociological makeup of the Fed (white male elites) can lead them to privilege the interests of the wealth holding class and ignore negative distributional effects on women and men of color and white women. Of course, the Fed's actions do not produce direct distributional effects; those are transmitted via the impact of interest rate changes on business and consumer borrowing, and as a result, on employers' decisions on whom to hire or fire in response to changes in demand.

In racially- and gender-equitable societies, race and gender differences in the probability of unemployment across business cycles would not exist, although individual probabilities of being unemployed might vary, stemming from differences in human capital and differential industry and occupation effects. Systematic intergroup differences in human capital and job concentration do exist, however, indicative of processes of group stratification that can explain at least a portion of race and gender differences in layoffs during downturns. A large body of gender and racial job competition research finds that women and men of color and women from dominant ethnic groups tend to be crowded into jobs and industries with low wages and benefits, characterized by employment volatility and absence of opportunities to move up the job ladder (Bonacich 1972; Hartmann 1976; Mason 1995, 1999; Standing 1989; Williams 1987, 1993; Williams and Kenison 1996). Job competition that slots subordinate groups for less stable jobs in lower-wage industries may indirectly contribute to differential gender and racial unemployment effects in response to interest rate hikes.

Several studies further suggest that overt discrimination is a cause of unequal unemployment rates by race and gender. For example, research on the cyclical patterns of employment finds that less than half the black-white male unemployment gap in the US can be attributed to observable factors other than race (Holcombe 1988; Stratton 1993; Sundstrom 1997). Similarly, Azmat, Guell, and Manning's (2004) investigation of female-male unemployment gaps in OECD countries fails to find support for human capital-related explanations. The authors did find, however, a correlation between gender gaps in unemployment and attitudes on men's deservingness of work when jobs are scarce, suggesting that hierarchical gender norms and stereotypes contribute to women's greater likelihood of experiencing unemployment during recessions.

Another body of research explicitly considers the impact of contractionary monetary policy by race and gender. Several studies, using VAR techniques, find that disinflationary policy (via higher federal funds rates) has unequal impacts on unemployment by race (Abell 1991; Thorbecke 2001; Carpenter and Rodgers 2004; Rodgers 2007). Based on impulse response functions, Thorbecke (2001) estimates that a one standard deviation increase in the nominal federal funds rate raises the difference between the black and white unemployment rate by 0.05 percentage points, compared to Carpenter and Rodgers's estimate of 0.15 percentage points. Thorbecke (2001) speculates that differentially negative impacts on blacks may be due to "ladder effects" wherein less skilled workers are laid off first, due to firm investment in training of higher skilled workers, or to a ratcheting upward of employers' selectivity, a less costly choice during recessions. Further, low-wage workers may also have less bargaining power in contrast to higher wage workers who are better able to protect their jobs during hard times. Another factor is discrimination in job access, likely to intensify in a labor market with job shortages as racial norms and stereotypes come into play in the job rationing process.

Evidence on the gendered impact of disinflationary policy is less consistent. Braunstein and Heintz (2008) find a negative impact on women's employment relative to men's in developing countries, using a method that examines outcomes following inflationary episodes. In contrast, Tachtamanova and Sierminska's (2009) recent study of OECD countries finds no evidence of systematic gender differences in unemployment rates. We contribute to this literature by assessing multiple discrimination in job loss due to monetary tightening and by evaluating whether structures of gender or of racial stratification dominate in situations of job scarcity induced by monetary tightening.

Stratification by Race and Gender: Complements, Substitutes, or Unrelated?

Racial Stratification: The Reproduction of Race Identity, Norms, and Stereotypes

To understand the interaction of race and gender hierarchies in labor markets, we consider here the emerging literatures on the economics of identity and stratification, which offer a framework for theorizing about how these hierarchies interact in labor markets in response to job shortages. We then integrate insights from the psychological and sociological literatures on prejudicial group attitudes.

A key theoretical argument is that racial identities are produced goods, responsive to shifts in the social and economic costs and benefits of holding such identities (Darity, Mason, and Stewart 2006). In the case of race (gender identity is discussed below), individuals sort along a continuum between two extreme identity formations, *racialized* and *individualist*. Racialists choose to identify with their own social group, and engage in collective action with those of similar identity to limit the outside group's control over material resources. They may do this explicitly by limiting job access, for example, or implicitly, by inculcating and perpetuating norms and stereotypes that shape perceptions of "deservingness." In contrast, individualists, as described by Darity, Mason, and Stewart (2006), have weak group identification and are willing to forgo status rewards that accrue to group conformity. Individualists eschew race identification as a means for assessing deservingness in access to and control over material resources. The share of the population that identifies as racialist or individualist responds to changes in material rewards for group identification.

We hypothesize that as the net benefit of group identification rises, the share of the population identifying as racialists will increase, with accentuated racialist norms that translate into discriminatory behavior in evidence. Macro-level influences may thus play an important role in attenuating or accentuating racialized behavior and racism by altering the costs and benefits of group identity. Jobs are a prized economic asset, and job scarcity is likely to accentuate the incentive of the dominant group to use racialized norms to improve their position in the job queue.

We might thus expect that during economic booms that produce broadly shared increases in income and employment opportunities, the share of racialists in the population will decline since the cost of holding an individualist identity decreases. In contrast, economic contractions may lead to an increase in the share of racialists in the population, palpably measured as a rise in job discrimination. Economic contractions or stagnation might be expected to lead to racial hysteresis effects, resulting in an increase in the share of racialists in the population as has emerged in Europe during the recent years of high unemployment and accentuated by the global crisis of 2008.[3]

Population density of the subordinate group may also act as a macro-level factor that determines the share of the dominant group with racialized identities. Holding constant other macro-level conditions (including rules on property ownership, legal consequences of discriminatory behavior, and so forth), the higher the population density of the subordinate group, the greater the perceived benefit to the dominant group of a racialized identity, which serves to limit competition over material resources.

Contact and *threat* theory offer hypotheses that describe special cases of the dominance of individualist or racialist identity norms.[4] Contact theory is associated with the work of Gordon Allport (1954), who held that race prejudice is an idiosyncratic individual attitude based on factually incorrect stereotypes, derived from the human propensity to categorize and summarize information. Contact theory's basic premise is that increases in intergroup contact, under structurally equitable conditions should lead to a revision of faulty stereotypes, reducing white prejudice against blacks.

Challenges to Allport's contact theory emerged early on. Herbert Blumer (1958) posited that race prejudice is not simply an individual state of mind, but rather, reflects a sense of group position. A feeling of superiority and hence a proprietary claim to privileges in certain areas undergird this prejudice. Following on Blumer's work, Blalock (1967) advanced a theory of group threat or competition (also called the visibility-discrimination hypothesis) to explain why racial inequality is higher in geographic areas with large concentrations of blacks. The latter approach stimulated research on the conditions under which whites perceive blacks a threat to white sense of group privilege.

Contact and threat theory reflect opposing predictions about the impact of interracial contact on the tendency to discriminate against blacks. A possible resolution to these apparently contradictory theories is explored in a number of studies that find "threshold" effects, with prejudice initially declining (increased contact causes whites to revise negative stereotypes) and then rising (the threat of competition is accentuated) as black population share rises (Forman 2003; Fossett and Kiecolt 1989; McCreary, England and Farkas 1989; Taylor 1998). Another body of evidence finds that the threat effect dominates at low percentages of blacks in the population, suggesting contact theory holds sway at higher black population shares. Pettigrew and Tropp's (2006) meta-analysis of intergroup contact theory studies leads them to conclude that although contact under a variety of conditions reduces prejudice, contact under unfavorable conditions may increase prejudice and tensions.

For the purposes of the current study, we posit that if threat effects undermine the benefits of contact under conditions of job scarcity, job losses resulting from disinflationary monetary policy will disproportionately affect blacks. Further, we can hypothesize that the ratio of black to white male unemployment rates will rise as the black share of the population increases. The results of previous studies imply that we may expect to find non-linearities in the relationship between monetary policy variables and race-based employment outcomes and that it is useful to examine multiple thresholds in this regard.

Gender and Threat Effects

Similar to racial identity formation, gender identities may fall along a continuum from *masculinist* to *gender egalitarian*. A masculinist identity reflects a patriarchal stance on gender relations, with adherents engaging in implicit or explicit collective action to ensure disproportionate economic and social power accrues to males (Braunstein 2008). *Gender egalitarians*, in contrast, adhere to norms that do not privilege one gender's resource control over another's.

Masculinists use their material and power advantage to maintain their preferential position in the construction of gender ideology, norms, and stereotypes that justify inequality (Blumberg 1984; Chafetz 1989). Increased resource scarcity might intensify the prevalence of masculinist identities among the population, leading to greater discrimination in job access. For the purposes of this study, we accordingly hypothesize that contractionary monetary policy leads to increases in the female to white male unemployment rate ratio. We further investigate the degree of stratification among women, assessing the relative effects of monetary tightening on black women's and white women's unemployment, both relative to white men.

How do race and gender stratification and inequality interact, particularly under conditions of job scarcity? Gender and racial discrimination may be complements, such that white women, black women, and black men all face relatively similar disadvantages in job access during economic downturns. Evidence of job competition between white women and all blacks suggests, however, that employers make hiring decisions based on a hierarchy of race/ethnic preferences (Reskin 1991; Mason 1999; Williams 1991; Spalter-Roth and Deitch 1999; Holzer and Neumark 2000; Stoll, Raphael, and Holzer 2004).[5] Waldinger (1997), for example, provides evidence that whites (including white women) are at the top of the job hierarchy, followed by Hispanics and blacks. In interviews with employers, Moss and Tilly (2001) also find a preference for hiring white women over other groups in labor markets where job skills have risen, with blacks perceived more negatively.

These studies suggest the plausibility of "nested" hierarchies of unemployment contingent on the degree of ethnic heterogeneity at the state level. More specifically, dominant groups (white men) may prefer to allocate joblessness to racially subordinate groups than to women of the dominant ethnic group. A rationale for this preference ranking is offered by a black supervisor in Button and Rienzo (2002: 16): "Hiring white women is a white man's way of making sure whites stay on top." White male racialists have a material incentive to shift the burden of joblessness to black men and women over white women, thus mitigating white family income losses. In short, when it comes to job discrimination, race invokes a greater penalty than gender.

The dynamics of race and gender stratification discussed here suggest three testable hypotheses. The first is that blacks and (all) women fare worse relative to white men when contractionary monetary decisions raise the real value of the policy interest rate, creating conditions of job scarcity. Second, to the extent threat effects influence outcomes, we hypothesize that the responsiveness of the black/white male unemployment rate ratio to policy interest rates varies with black population density. Third, we explore the possibility that racial hierarchies dominate gender hierarchies by assessing the impact of black share of the population on the responsiveness of the ratio of female to white male unemployment rates to changes in interest rate policy.

Empirical Analysis

The Modeling Approach

The empirical model we construct has the primary goal of assessing the distributional impact of interest rate policy on black men and women and white women relative to white men. We test subordinate group effects by gender and race separately, with the ratio of female and black unemployment rates, relative to white male unemployment rates, respectively, serving as our dependent variables. We then disaggregate subordinate group effects, using black female, black male, and white female relative unemployment rates as dependent variables. Employing a panel data set of US states, we are able to take into account fixed effects, that is, unobserved state-level differences that influence outcomes.

We focus the analysis on a primary monetary policy instrument used by the Federal Reserve, the federal funds rate, measured as the interest rate on overnight loans between banks. The Federal Reserve attempts to influence macroeconomic outcomes by raising and lowering the federal funds rate in response to changes in inflation, economic performance, and employment. The real federal funds rate impacts unemployment by influencing the macroeconomic performance of the US economy as a whole. However, our panel data are disaggregated to the state level. Therefore, one challenge is to distinguish the impacts of macroeconomic policies that operate at the national level from regional economic dynamics that may operate independently of national policy and which vary from state to state.

Data

We assembled a panel dataset for each of the 50 states covering the period 1979 to 2008 using four sources: the Current Population Survey (CPS), the Bureau of Economic Analysis (BEA), the Federal Reserve Board of Governors, and the Bureau of Labor Statistics (BLS).[6] Annual labor market statistics, including state-level disaggregated estimates of employment, unemployment, and labor force participation by race, gender, and ethnicity, were calculated directly from the CPS micro-data. The BEA produces state and national-level estimates of GDP. The

Federal Reserve was our source for the federal funds rate, and the BLS maintains the consumer price index, used to calculate annual inflation rates.

Merged CPS data on the outgoing rotation group were used to estimate the annual state-level labor market statistics, disaggregated by race/ethnicity and gender.[7] We apply the methodology developed by the Center for Economic and Policy Research (CEPR) to classify individuals into four mutually exclusive racial/ethnic groups: white, black, Hispanic, and other. Because of the small sample size in the out-going rotation group, reliable estimates of the unemployment rate for blacks were not possible in states with very low black shares of the state population. Since construction of our dependent variable requires an estimate of the black unemployment rate over time, we dropped states with more than 10 missing observations due to excessively small samples. This resulted in 12 states being dropped, all of which have very small black population shares: Hawaii, Idaho, Maine, Montana, New Hampshire, New Mexico, North Dakota, Oregon, South Dakota, Utah, Vermont, and Wyoming.

To distinguish macroeconomic dynamics that affect aggregate output at the national level from state-specific changes in economic activity, we regressed state-level GDP growth on national-level GDP growth using a fixed-effects model. We then captured the residuals (both the random errors and the fixed effects components of the error term) and used these residuals as an indicator of state-level changes in real economic activity, removing the impact of variations at the national level.

Times series data are potentially non-stationary, which, if not addressed, can produce spurious results. We therefore conducted panel unit root tests for all variables. Discussion of the methodology used is provided in the appendix, with test results summarized in Table A.1. We rejected the presence of a unit root in all cases.

Analysis

We use the panel dataset to estimate the following relationship:

$$U_{it}^{z} = \beta_0 + \beta_1 FFR_t + \beta_2 LFPR_{it}^{z} + \beta_3 Gr_{it} + \beta_4 BLSH_{it} + \beta_5 BLSH_{it}^{2} + \eta_i + \varepsilon_{it} \tag{1}$$

where U^z is the ratio of the subordinate group unemployment rate to the white male unemployment rate; the subscripts i and t index states and years, respectively; *FFR* is the real federal funds rate (the nominal rate less the inflation rate);[8] $LFPR^z$ is the labor force participation rate ratio of the subordinate group to white males; *Gr* is the state-level growth of output after the impact of national-level growth dynamics has been removed; *BLSH* and $BLSH^2$ are the black share of the population and black share squared, respectively; η is the component of the disturbance term associated with state-specific effects; and ε is a random error term.

It is useful to consider potential endogeneity of two variables, the federal funds rate and relative labor force participation rates. With regard to the former, we deem endogeneity concerns to be negligible. The Federal Reserve is unlikely to propose national interest rate adjustments in response to state-level changes in the unemployment rate ratio, given the degree of heterogeneity among the states.

In contrast, labor force participation rates may vary inversely with unemployment, capturing the "discouraged worker" effect.[9] Our motivation for including labor force participation as an explanatory variable is to correct a potential bias with unemployment rates as conventionally measured. If high unemployment reduces labor force participation, standard unemployment rates underestimate the effect of monetary policy because lower labor force participation reduces measured unemployment. In that sense, our regression results produce a lower-bound estimate of unemployment effects. More succinctly, the labor force participation variable addresses an issue about the measurement of unemployment; it is not a direction of causality issue.

Equation (1) directly incorporates the black population share as an explanatory variable. Since the squared population share is also included, the relationship is non-linear. This represents one strategy for modeling non-linearities in terms of the unemployment rate ratios. However, the coefficients on the other variables remain constant with variations in the black population share. An alternative approach to capturing non-linearities in the responsiveness of relative unemployment rates to monetary tightening is to develop threshold models in which the coefficients themselves are allowed to vary when the black population share falls above or below certain thresholds.

Therefore, we test for threshold effects on unemployment rate ratios, anticipating that the estimated coefficients will vary depending on the black share of the working age population. However, we treat the thresholds at which the structure of the relationships changes as unknown. Therefore, as a first step, we estimate the thresholds of the black population share at which the relationship between the federal funds rate and differential race and gender outcomes changes. We then generate different estimates of the model for states whose black population shares fall above or below particular thresholds.

To maintain a minimum number of observations, we additionally require that any division based on the threshold retain at least 4 states, placing an upper limit on our thresholds of approximately 28 percent. In only four states does the black share exceed 27 percent: Georgia, Louisiana, Mississippi, and South Carolina. For the one-threshold model, we estimate a series of equations, allowing the threshold, τ, to vary from a low of one percent to a high of 28 percent. For each value of τ, we estimate two equations—one for all states whose average black population share falls below τ and one for states whose population share is greater than or equal to τ. The set of estimates with the highest regression sum of squares is taken as the best fit and determines the value of τ we use in this analysis.

A similar procedure is used in the two-threshold model, except that we have two unknown thresholds, τ_1 and τ_2. We allow τ_1 to vary between one percent and 28 percent—again, imposing the requirement that each sub-group of states contain at least four states. For each value of τ_1, we allow τ_2 to vary throughout a similar range as long as $\tau_2 > \tau_1$ for any given τ_1. For each value of τ_1 and τ_2, we estimate three equations: one for states whose black population share falls below τ_1, a second for states whose black population share is greater than τ_1 but less than τ_2, and a third for states whose black population share is greater than or equal to τ_2. Our point estimates of the values of τ_1 and τ_2 are those that maximize the total regression sum of squares.

For the models with the black/white male unemployment rate ratio as the dependent variable, we find that the two-threshold model has the best fit. The thresholds that maximize the sum of squares of the regression are 11 percent and 25 percent. In the case of the one-threshold model, the threshold value that maximizes the regression sum of squares is 25 percent. Table 1 reports these results and compares the regression sum of squares of the two-threshold model with those of the one-threshold model and of a model with no thresholds imposed.

The two-threshold model provides the best fit for the gender models, with the female/white male unemployment rate ratio as the dependent variable. The thresholds that maximize the sum of squares of the regression are 14 percent and 25 percent. These thresholds are close to those from the black/white male model. This suggests that structural changes that are associated with different black population shares affect the estimated relationships in both the race and gender models at similar threshold levels. For comparative purposes, we also estimate a two-threshold model for the female/white male relationship using the same thresholds from the black/white male unemployment rate regressions (11 and 25 percent). The regression sum of squares is slightly lower than when the thresholds are set at 14 percent and 25 percent, but the difference is negligible. These results are summarized in Table 1.

Hansen (1999) presents a method for formally testing the existence of threshold effects in panel data. The null hypothesis of no threshold effects is tested against the alternative of the existence of threshold effects. The test statistic is a likelihood ratio test based on an F-statistic calculated from the residual sum of squares under the null and alternative hypotheses. Since the thresholds are themselves estimated and do not apply under the null hypothesis, the distribution of the test statistic is irregular and critical values cannot be tabulated. We employ a bootstrap procedure recommended by Hansen (1999) for estimating the distribution and obtaining p-values which are asymptotically valid. Table 1 presents the test statistic and estimated p-values of this test of the existence of one- and two-thresholds against the null hypothesis of no threshold based on 1,000 bootstrap iterations. In all cases, we are able to reject the null hypothesis of no threshold effects.[10]

Note that we did not test the model in which the 11 percent threshold was applied to the gender estimations since the point estimate of the threshold is 14 percent. Hansen (1999) describes a procedure for estimating confidence intervals around the estimated thresholds, which are random variables. For the lower threshold of the

Table 1

Estimates of black population share thresholds, total regression sum of squares

Estimation	Regression sum of squares	Bootstrap threshold test results
	Black/white male unemployment	
No threshold	10.42	n/a
One threshold (25%)	15.91	8.97
		(p-value = 0.014)
Two thresholds (11% and 25%)	21.62	18.46
		(p-value < 0.001)
	Female/white male unemployment	
No threshold	2.77	n/a
One threshold (25%)	4.94	32.78
		(p-value = 0.021)
Two thresholds (14% and 25%)	6.32	54.70
		(p-value = 0.007)
Two thresholds (11% and 25%)	6.28	—

Note: p-values are achieved significance levels based on 1,000 bootstrap repetitions.

two-threshold gender model, the 95 percent confidence interval is 8 to 15 percent range, which includes the alternative threshold of 11 percent.

Table 2 presents the detailed coefficient estimates of the basic models and the two-threshold fixed effects models with thresholds of 11 percent and 25 percent for the black/white male regressions and 14 percent and 25 percent for the female/white male regressions. Columns 1 and 5 present estimates of the model without threshold effects as described in Equation (1) for relative unemployment rates of blacks and females, respectively. Columns 2–4 and 6–8 give results of the two-threshold models with black and female relative unemployment rate ratios as the dependent variables, respectively (and with the black population share omitted as an explanatory variable since it is used to determine the relevant thresholds).

Table 2

Estimated coefficients of two-threshold model, fixed effects

	Black/white male unemployment rate ratio				Female/white male unemployment rate ratio			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Explanatory variables	Full sample	<11%	11% to 25%	>25%	Full sample	<14%	14% to 25%	>25%
Constant	2.291	3.045	3.847	−2.179	0.245	0.318	2.798	−1.826
	(0.61)*	(0.65)*	(0.60)*	(2.41)	(0.31)	(0.15)*	(0.58)*	(1.45)
Federal funds rate	0.031	0.009	0.048	0.071	0.024	0.019	0.002	0.086
	(0.01)*	(0.01)	(0.01)*	(0.02)*	(0.01)*	(0.003)*	(0.01)	(0.02)*
Labor force participation	−0.610	−0.661	−1.869	5.900	0.695	0.901	1.776	4.169
rate	(0.62)	(0.75)	(0.67)*	(2.82)	(0.40)	(0.12)*	(0.68)*	(1.77)
State growth	−0.701	−0.021	3.852	−0.304	0.131	−0.143	1.695	1.076
	(0.70)	(0.78)	(1.21)*	(0.708)	(0.24)	(0.21)	(0.45)*	(0.27)*
Black share	0.076				0.037			
	(0.06)				(0.02)*			
Black share squared	−0.001				−0.0004			
	(0.001)				(0.0004)			
Number of states	38	21	13	4	38	27	7	4
N	1,088	595	377	116	1,102	783	203	116
Regression sum of squares	12.711	2.367	14.637	5.183	3.337	1.280	1.627	3.590

Note: Robust standard errors in parentheses. * indicates p-value less than or equal to 5%.

Consider first the estimates of the determinants of the black/white male unemployment rate ratio. Column 1 shows the estimated coefficients for all states included in the panel. The constant term is 2.291, consistent with past research indicating the black/white unemployment rate ratio hovers around 2. A one percentage point increase in the real federal funds rate raises the black/white male unemployment rate ratio by 0.031 percentage points. Racial differences in labor force participation do not have a significant effect on the dependent variable, nor does the adjusted state growth rate. Neither the black share of the population nor its square is statistically significant. The estimated coefficients for states with a black population share of less than 11 percent are given in Column 2. In this group of states, none of the coefficients are statistically significant, with the exception of the constant term. In states whose black population share lies between 11 and 25 percent (column 3), the real federal funds rate exerts a positive significant effect on the ratio of black to white male unemployment. This relationship becomes even stronger when the black population share exceeds 25 percent (column 4), with a one percentage point increase in the federal funds rate raising the black/white male unemployment rate ratio by 0.071 percentage points. These results are consistent with threat theory, whereby exclusion and discrimination against blacks increases with rising black population density.[11]

The coefficient on state-level growth dynamics (controlling for national-level growth) is positive and statistically significant in states with a black population share between 11 and 25 percent. Faster regional growth, controlling for national-level growth, raises the ratio of black to white male unemployment, indicating blacks experience a disproportionately smaller boost to employment from regional sources of growth in states with a black population share between 11 and 25 percent. Noting that the coefficient on this variable for the other threshold groups is negative, but not statistically significant, these results suggest that racially based job exclusion does indeed depend on black population density.

Turning to the female/white male results for all states for which we have sufficient data on black unemployment (column 5), the coefficient on the real federal funds rate is 0.024, smaller than the coefficient in the corresponding black/white male regression and statistically significant. The only other significant variable in that regression is the black share of the population. It is positive, indicating that women's unemployment rate relative to white males rises linearly with increases in the black share of the population. This is consistent with a hypothesis of job competition, potentially due to the crowding of subordinate groups into a limited number of job slots, relative to white men.

In the threshold equations with the female/white male unemployment rate ratio as the dependent variable (columns 6–8), increases in the federal funds rate exert a positive and statistically significant effect when the black population share is below 14 percent or above 25 percent. In the middle range, this coefficient is not statistically different from zero. For states with less than 14 percent black population share, the coefficient value is 0.019. This effectively drops to zero for states with a black population share between 14 percent and 25 percent. However, for states with average black population shares greater than 25 percent, the size of this coefficient increases to 0.086. In states with black population shares in excess of 14 percent, state-level growth exerts a positive and statistically significant impact on the ratio of female to white male unemployment. Thus, similar to the impact of economic growth on blacks, women seem to be last hired during economic upturns, at least in states with a black population share greater than 14 percent.

We also estimated two variations on the female/white male unemployment rate ratio equations. First, we impose the same thresholds (11 and 25 percent) as we applied to the black/white male unemployment rate equations in Table 2 in order to facilitate race/gender comparisons. We found that the coefficient estimates on the federal funds rate variable behave in a similar fashion to those in Table 2. Second, we generated coefficient estimates including all 50 states. Recall that we dropped states from the sample because it was not possible to estimate black unemployment rates for states with very small black populations, but this constraint does not apply to women's unemployment rates. Using the same methodology and including all 50 states, we find that the optimal thresholds are again 14 percent and 25 percent. Including the full set of states only alters the coefficient estimates for states with less than a 14 percent black population share. The coefficient estimates for the sample using states with less than a

14 percent black population share are quite similar to those in Table 2 generated from a more limited sample. The full set of results of these variations are available on request.

Robustness Tests: Controlling for Education and Employment Concentration

Gender and racial differences in unemployment may be due to processes that reflect gender and racial stratification in education and job segregation. For that reason, we carry out a robustness check, controlling for gender and racial differences in the share of respective populations with a college education and the relative shares employment in interest rate-sensitive industries. We identified construction and durable goods manufacturing as the primary interest-rate sensitive industries, following Thorbecke (1997). State-level gender and race data on education and employment by industry are from the CPS, as discussed above.

The ratios of the percentage of the labor force with some college/tertiary education by race and gender are measured respectively as:

$$COLL^{BWM} = \frac{\%COLL^{B}}{\%COLL^{WM}}, \; COLL^{FWM} = \frac{\%COLL^{F}}{\%COLL^{WM}}$$

where percent *COLL* is the percentage of a group's labor force participants with some college education, even if they did not earn a degree, and *BWM* (*FWM*) is the ratio of blacks to white males (females to white males). Using similar notation, the percentage of blacks (females) employed in interest-rate sensitive industries, relative to the white male share is:

$$\%IND^{BWM} = \frac{\%IND^{B}}{\%IND^{WM}}, \; \%IND^{FWM} = \frac{\%IND^{F}}{\%IND^{WM}}$$

where *IND* denotes the share of the respective groups employed in interest-rate sensitive industries, and the remaining terms are defined as for education.

We expect a negative coefficient on the percentage of blacks (females) relative to white males with a college education if there are "ladder" effects in job losses during recessions whereby less skilled workers are the first to be laid off (Jefferson 2005). Conversely, as the share of employed blacks (females) working in interest-rate sensitive industries rises relative to white male concentration in these industries, we anticipate an increase in the corresponding unemployment rate ratio. This captures the combined effects of job concentration and employment in industries sensitive to increases in borrowing costs. In addition to providing the means to conduct a robustness check on the federal funds rate variable, inclusion of the additional variables allows us to parse the mechanisms of stratification and employment disadvantage by race and gender into three component parts: discrimination in job access, educational inequality, and job concentration (or segregation).

Consider first the estimates of black/white male unemployment rate ratios in Table 3. Column 1 shows the coefficients for all states for which there are sufficient data. Coefficients on education and concentration in interest rate-sensitive industries are statistically significant and, as expected, work in opposite directions. The black/white proportion of college-educated workers has a negative effect on the unemployment rate ratios, suggesting that part of the raw unemployment gap is explained by white males' greater probability of having a college education. The coefficient on the employment concentration variable indicates that as the relative share of blacks employed in interest rate-sensitive industries rises, the black/white male unemployment rate ratio increases. Controlling for these two variables, the federal funds rate continues to exert a positive significant effect on the unemployment rate ratio, and is somewhat smaller in magnitude than in the regressions that do not control for education and job segregation (0.022 as compared to 0.031).

The threshold model results (columns 2–4) also indicate that the inclusion of education and employment variables does not alter the significance of the coefficients on the federal funds rate, labor force participation, and state growth in the analogous models in Table 2 although the size of federal funds rate coefficients declines slightly. The education variable is only significant in the states where blacks comprise 11 to 25 percent of the population, while the job segregation variable is not significant in any of the threshold models.

Table 3

Estimated coefficients, controlling for education and job segregation

	Black/white male unemployment rate ratio				Female/white male unemployment rate ratio			
Explanatory variables	(1) Full sample	(2) <11%	(3) 11% to 25%	(4) >25%	(5) Full sample	(6) <14%	(7) 14% to 25%	(8) >25%
Constant	1.583	2.332	2.453	–3.569	0.047	0.089	2.715	–1.48
	(0.98)	(0.85)*	(1.25)	(2.06)	(0.34)	(0.15)	(0.72)*	(1.37)
Federal funds rate	0.022	0.002	0.035	0.057	0.022	0.017	–0.0002	0.070
	(0.01)*	(0.02)	(0.01)*	(0.01)*	(0.01)*	(0.01)*	(0.01)	(0.02)*
Labor force participation rate	1.330	–0.109	0.531	9.945	1.176	1.127	–1.184	6.009
	(0.96)	(0.94)	(1.25)	(3.14)*	(0.50)*	(0.27)*	(1.13)	(1.26)*
State growth	2.064	2.772	3.545	0.141	0.105	–0.107	1.773	0.616
	(1.24)	(2.02)	(1.80)	(1.43)	(0.25)	(0.23)	(0.45)*	(1.26)
Black share	0.023				0.043			
	(0.07)				(0.02)*			
Black share squared	0.0001				–0.0005			
	(0.002)				(0.0004)			
College education	–1.560	–0.860	–1.458	–2.678	–0.185	–0.040	–0.237	–1.067
	(0.43)*	(0.84)	(0.37)*	(1.04)	–0.09	(0.09)	(0.19)	(0.22)*
Industry	0.400	0.579	0.421	–0.161	0.398	0.446	0.144	0.455
	(0.19)*	(0.29)	(0.22)	(0.58)	(0.18)*	(0.17)*	(0.92)	(0.64)
Number of states	34	17	13	4	38	27	7	4
N	470	144	245	81	1,102	783	203	116
Regression sum of squares	17.093	4.124	13.041	7.389	4.221	1.567	1.627	5.268

Note: Robust standard errors in parentheses. * indicates p-value less than or equal to 5%.

The results from the gender regressions (columns 5–8) produce similar results to those for blacks/white males. The higher the ratio of females to white males with a college education, the lower the unemployment rate ratio while employment concentration of women in interest-rate sensitive industries relative to white men raises the ratio. The federal funds rate coefficients are slightly lower than in the restricted regressions (columns 5–8 in Table 2), but retain their statistical significance.

Black Women, White Women, and Black Men: Is There a Hierarchy Within the Subordinate Groups?

A challenge in assessing the role of gender and race as categories of stratification is that they overlap. Conceivably, our results on the harmful effect of disinflationary monetary policy on all women's relative job prospects could be capturing a differentially negative effect on black women. In an effort to further refine our understanding of stratification dynamics and the role of multiple discrimination, we re-run the regressions separately for black and white women and black men, all relative to white men.

Results are presented in Table 4. (Table A.2 in the appendix provides results of the robustness check where basic regressions are augmented with controls for education and employment.) Columns 1, 5, and 9 in Table 4 present estimates of the model without threshold effects for black women, white women, and black men, respectively. A higher federal funds rate has a positive significant effect on the ratio of black female, white female, and black male unemployment relative to white males. The size of the effect, however, differs systematically among the three subordinate groups. The impact of a one percentage point increase in the federal funds rate on white women's relative unemployment rate is 0.015, compared to 0.028 for black men, and 0.043 for black women. The interest rate effect on black women's relative unemployment increases with black population density.

The effect on white women's relative unemployment, in contrast, is lower in states with black population density between 11 and 25 percent, with a statistically significant effect in states with black population density in excess of 25 percent. Black men's relative

Table 4

Estimates of the determinants of subordinate groups' unemployment rates relative to white men's

	Black female/white male unemployment rate ratio				White female/white male unemployment rate ratio				Black male/white male unemployment rate ratio			
Explanatory variables	(1) Full sample	(2) 11%	(3) 11% to 25%	(4) >25%	(5) Full sample	(6) 11%	(7) 11% to 25%	(8) >25%	(9) Full sample	(10) 11%	(11) 11% to 25%	(12) >25%
Constant	0.834	1.615	3.653	0.051	0.497	0.294	1.478	−0.725	3.716	4.274	1.899	0.249
	(0.636)	(0.42)*	(0.57)*	(1.79)	(0.19)*	(0.17)	(0.22)*	(0.67)	(0.92)*	(0.84)*	(0.86)*	(1.54)
Federal funds rate	0.043	0.007	0.052	0.113	0.019	0.0179	0.008	0.049	0.028	0.027	0.043	−0.022
	(0.01)*	(0.01)	(0.01)*	(0.04)*	(0.004)*	(0.003)*	(0.003)*	(0.02)	(0.01)*	(0.01)*	(0.006)*	(0.03)
Labor force part. rate	0.442	0.977	−1.887	3.097	0.526	0.680	0.977	2.147	−1.447	−1.948	−0.285	3.33
	(0.48)	(0.49)*	(0.64)*	(2.07)	(0.23)	(0.21)*	(0.50)	(0.77)	(0.76)	(0.89)*	(0.27)	(1.71)
State growth	0.766	0.033	3.935	1.543	−0.211	−0.071	0.033	0.649	0.555	0.004	3.244	−1.942
	(0.59)	(0.65)	(1.56)*	(1.27)	(0.18)	(0.22)	(0.65)	(0.40)	(0.93)	(1.13)	(1.11)*	(1.16)
Black share	0.151				−0.004				−0.008			
	(0.07)*				(0.02)				(0.09)			
Black share squared	−0.003				0.0001				0.001			
	(0.001)*				(0.003)				(0.002)			
Number of states	38	21	13	4	38	21	13	4	38	21	13	4
N	1,095	602	377		1,102	609	377	116	1,095	602	377	116
Regression sum of squares	0.834	4.308	24.201	9.111	2.466	0.974	1.25	1.25	18.849	19.118	6.73	2.304

Note: Robust standard errors in parentheses. * indicates p-value less than or equal to 5%.

unemployment rate rises in response to a one percentage point increase in the real higher federal funds rate from 0.027 in states with black population density below 11 percent to 0.043 in states with a density between 11 to 25 percent. Interest rate effects on black men's relative unemployment are not statistically significant in states with black population shares over 25 percent.

State growth and labor force participation rate ratios do not exert a statistically significant influence on unemployment rate ratios for any of the subordinate groups in the no-threshold models, but raise relative black female and black male rates in states with a black population density of 11 to 25 percent. State growth does not significantly affect white women's relative unemployment rate at any level of black population density.

Disaggregation by race and gender allows us to conclude that blacks disproportionately bear the costs of unemployment induced by disinflationary monetary policy relative to white women, and that burden is heaviest for black women. The data indicate that the penalty for black women and men is positively correlated with black population density, consistent with threat theory. Figure 1 provides a visual representation of the differential impacts on black women and black men relative to white women as the federal funds rate rises, based on results in Tables 4 and A.2.

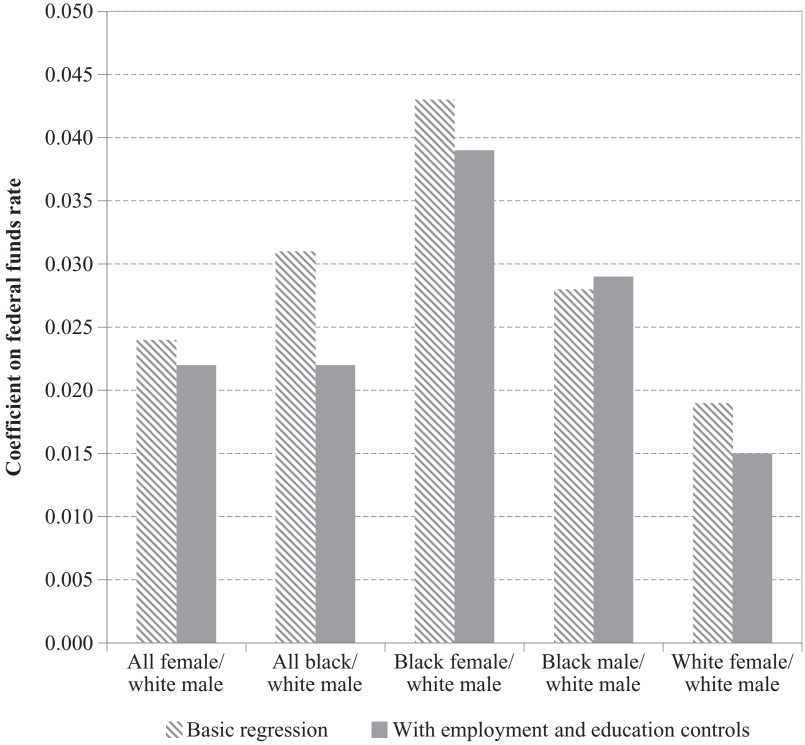
Race and Gender Stratification: Substitutes or Complements?

We posited that there may be a relationship between black and female relative unemployment rate ratios, and in particular that female/male rate ratios might fall as black population share rises. This would suggest that in predominantly white states, (white) women would bear a disproportionate share of unemployment resulting from interest rate hikes, relative to white men. But as the black share of the population rises, we hypothesized that the job costs of disinflationary monetary policy would be shifted to blacks, consistent with threat theory.

Figure 2 is constructed from the data in Table 5 to compare black female, black male, and white female interest rate effects as black population share rises. We score coefficients with a p-value > 0.05 as

Figure 1

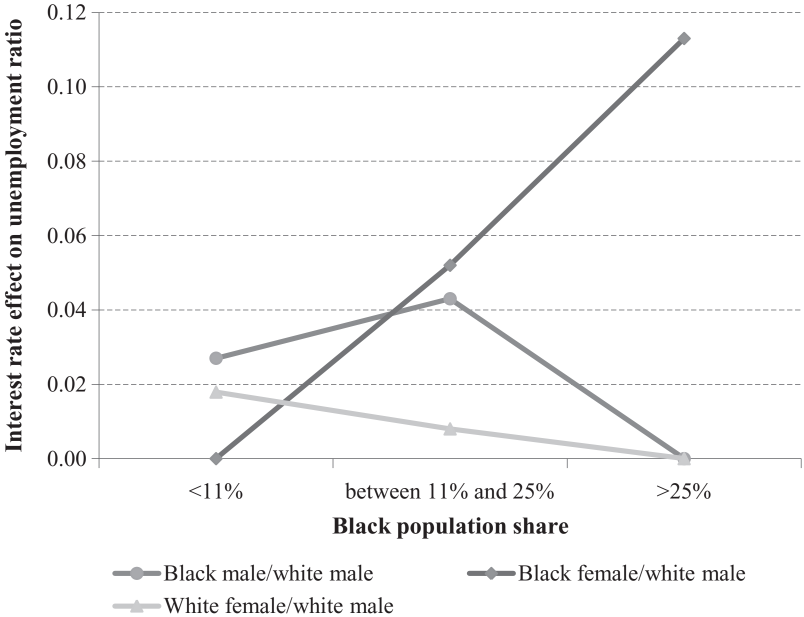
Impact of one unit increase in real federal fund rates on black and female unemployment rate ratios, disaggregated



0. At black population shares below 11 percent, the effect of a hike in the federal funds rate on the unemployment rate ratio of the subordinate group to white males suggests the following hierarchy: white males and black females, followed by white females and then black males. However, in states with black population shares ranging from 11 to 25 percent, black women especially and then black men are substantially more likely than white women to be put at the back of the job queue, all relative to white men, when contractionary monetary policy is pursued. At black population density above 25 percent, the interaction of race and gender effects is in evidence with strong

Figure 2

Disaggregated comparison of interest rate effects on unemployment rate ratios by black share of population



effects on black women's relative unemployment but no significant relative effects on white women or black men.

We note, however, that only four states have black population shares in excess of 25 percent, and therefore focus on the change in interest rate coefficients as population share rises from below 11 percent to the group with shares between 11 and 25 percent. The results for this group of states suggest that racialized identity norms dominate, consistent with the threat hypothesis. Thus, at low black population shares, white men, whether as employers or workers able to influence hiring and firing decisions, shift the burden of monetary policy-induced unemployment to white women and black men. But as the black population share rises, the burden of unemployment shifts heavily away from white women and toward blacks, both male and female. That is, racialized norms that give whites preferential access to jobs appear to dominate over gender norms which infer that men are more deserving, when jobs are scarce.

Conclusions

We can conclude from this analysis that the effects of monetary tightening are neither race- nor gender-neutral. The impact weighs heavily on black men and women, and white women, with a significantly greater penalty for being black, whether one is male or female. Racial and gender differences in college education and job concentration in interest rate-sensitive sectors do not explain away the differentially negative impact of monetary policy on these subordinate groups. The results presented here confirm that gender analysis in the context of an ethnically heterogeneous society such as the US requires attention to potentially differential effects by ethnicity that may be stronger than gender differences.

Another implication of our results is that the distributional effects of the Federal Reserve's monetary policies should inform their decision-making. Given the long-term impact of unemployment on adults and their children (Darity and Goldsmith 1996), we might indeed be concerned about whether monetary policy contributes to the reproduction of poverty and inequality between whites and blacks, and women and men, in particular women who are lone mothers. The long-run negative effects of inequality have been established in a variety of studies. The Federal Reserve's failure to note the distributional consequences of its policy actions may in fact contribute to long-run inflationary pressures, resulting from the slowdown in labor productivity growth that inequality produces.

Notes

1. In contrast to perceived inflation, monetary policy responses to recessions may be more varied than lower interest rates as the current period of quantitative easing in the US shows.

2. Limited state-level data on Hispanics and coding ambiguities led us to focus exclusively on black-white unemployment gaps. Hispanics are treated as an ethnic rather than racial group in the Census. Further, emerging research

links disadvantage in labor markets to persons with darker skin shades, suggesting that race/ethnicity is not a dichotomous category (Goldsmith, Darity, and Hamilton 2007). Frank, Akresh, and Bu (2010) found that 79 percent of Latino respondents in the New Immigrant Survey identified themselves as white, regardless of their skin color. Inclusion of Hispanics as a distinct ethnic group in this study therefore could confound results. That said, the absence of adequate Hispanic data leaves a lacuna, given increasingly complex skin color and racial/ethnic signals that affect labor market outcomes. Previous studies on race and gender earnings and job displacement effects find that Hispanic women rank below white and then black women, and Hispanic men below white men and above black men (Spalter-Roth and Deitch 1999). Our study therefore offers only a partial race-gender ranking.

3. The increase in unemployment, combined with ominous discussions of tax increases and budget cuts, bolstered the political right in European elections in 2009. Election campaigns, marked by anti-immigrant messages linked to job shortages, resulted in right-wing parties making electoral gains in the Netherlands, Italy, Hungary, Great Britain, and Austria (Margaronis 2009).

4. We are grateful to Patrick Mason for this observation (private communication, June 15, 2010).

5. For a discussion of multiple discrimination, see Brewer, Conrad, and King (2002) and Ruwanpura (2008).

6. Given the unique structure of the Washington, DC, economy, we treat it as on outlier and do not include it in this analysis.

7. We are indebted to John Schmidt of the Center for Economic and Policy Research for his expertise in developing these estimates.

8. Because unemployment effects of contractionary monetary policy have been found to peak at five quarters (Christiano, Eichenbaum, and Evans 1996), we also ran regressions with current and lagged values of the real federal funds. The sum of the coefficients on current and lagged values of the federal funds rate was comparable to the coefficients in our models with only the current federal funds rate. Results available from the authors on request.

9. While the labor force participation variable is potentially endogenous, Carpenter and Rodgers (2004) found that increases in the federal funds rate lower the employment-population ratio of minorities by raising unemployment, not by lowering labor force participation rates.

10. Our empirical model differs from Hansen's (1999) in one respect—our thresholds are time-invariant. Nevertheless, we applied the same general procedure of creating a bootstrap sample from estimated residuals, assuming that the null hypothesis is valid. We also tested the hypothesis of a two-threshold model against the null of a one-threshold model. For both the race and the gender estimations, we were able to reject the null hypoth-

esis at a 1-percent significance level (for the race estimates, the estimated p-value was 0.008; for the gender equations the p-value was less than 0.001).

11. Comparisons of point estimates with previous studies are not meaningful due to differences in methodology and measurement of the relationship between black and white unemployment rates. We use white males as a measure of the dominant group, and the unemployment rate ratio, a non-linear indicator. Thorbecke (2001) and Carpenter and Rodgers (2004)'s dependent variable is the difference between black and (all) white unemployment rates, a linear measure.

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Appendix

Panel Unit Root Tests

We tested all the variables in our panel for unit roots using Fisher-type panel unit root tests with an augmented Dickey-Fuller specification applied to the individual cross-sections. We used the Fisher test because our panel is slightly unbalanced due to occasional missing observations for certain race-disaggregated variables. Other unit root tests (for example, Im, Persaran, and Shin) require precisely balanced panels.

Table A.1 summarizes the results of the Fisher panel unit root tests. The Fisher test assumes an AR(1) process in the specification of the underlying Dickey-Fuller specification. Columns (1) and (2) of Table A.1 report the results of the basic Fisher test, with Column (2) incorporating a deterministic time trend. Columns (3) and (4) augment the basic Fisher test with an additional lagged difference term, with Column (4) including a deterministic time trend.

The results of the test show that the vast majority of variables are stationary (that is, have no unit root) across the different specifica-

Table A1

Fisher-type panel unit root tests, χ-squared test statistics (p-values in parentheses)

	(1)	(2)	(3)	(4)
U^{FW}	733.9	618.1	341.1	272.4
	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)
U^{BWM}	663.0	576.6	348.1	285.3
	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)
$LFPR^{FWM}$	88.1	113.2	107.1	64.2
	($p = 0.16$)	($p = 0.004$)	($p = 0.01$)	($p = 0.83$)
$LFPR^{BWM}$	242.8	316.5	168.4	213.5
	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)
Gr	388.3	293.4	337.3	213.9
	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)
BLSH	136.4	166.3	120.4	161.5
	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)
$COLL^{FWM}$	71.6	296.3	62.8	293.8
	($p = 0.62$)	($p < 0.001$)	($p = 0.86$)	($p < 0.001$)
$COLL^{BWM}$	225.4	260.1	165.5	228.1
	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)
IND^{FWM}	286.4	515.9	155.9	314.7
	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)
IND^{BWM}	51.0	80.1	63.8	148.2
	($p = 0.43$)	($p = 0.004$)	($p = 0.01$)	($p < 0.001$)
U^{BFWM}	677.7	587.8	301.1	250.0
	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)
U^{WFWM}	849.6	748.3	429.9	376.0
	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)
U^{BMWM}	785.9	675.7	429.1	358.1
	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)
$LFPR^{BFWM}$	225.0	311.8	168.9	216.6
	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)
$LFPR^{WFWM}$	96.2	107.6	99.3	65.6
	($p = 0.06$)	($p = 0.01$)	($p = 0.04$)	($p = 0.80$)
$LFPR^{BMWM}$	498.1	439.0	329.9	292.5
	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)	($p < 0.001$)

Note: Column 1: AR(1) process, no deterministic trend. Column 2: AR(1) process, deterministic trend. Column 3: AR(1) process with additional lagged difference term, no deterministic trend. Column 4: AR(1) process with additional lagged difference term, deterministic trend.

Table A2

Black women and men and white women's unemployment, relative to white men, robustness check

Explanatory variables	Black female/white male unemployment rate ratio				White female/white male unemployment rate ratio				Black male/white male unemployment rate ratio			
	(1) Full sample	(2) 11%	(3) 11% to 25%	(4) >25%	(5) Full sample	(6) 11%	(7) 11% to 25%	(8) >25%	(9) Full sample	(10) 11%	(11) 11% to 25%	(12) >25%
Constant	1.846	2.352	3.04	−0.183	0.375	0.037	1.508	−0.359	2.639	3.249	2.326	−0.171
	(1.15)	(1.08)*	(1.13)*	(1.75)	(0.22)	(0.19)	(0.25)*	(0.82)	(1.04)*	(0.85)*	(0.85)*	(2.16)
Federal funds rate	0.039	0.020	0.052	0.068	0.015	0.014	0.006	0.041	0.029	0.038	0.016	−0.029
	(0.01)*	(0.22)	(0.02)*	(0.01)*	(0.004)*	(0.003)*	(0.003)*	(0.23)	(0.01)*	(0.02)*	(0.01)	(0.04)
Labor force	1.468	1.143	−0.254	6.72	1.082	1.023	−0.041	3.111	0.136	−0.285	1.118	3.606
part. rate	(0.99)	(1.17)	(1.30)	(2.27)*	(0.32)*	(0.32)	(0.47)	(1.00)*	(0.59)	(0.71)	(1.02)	(1.52)
State growth	2.065	2.424	3.3	0.091	−0.113	−0.111	−0.147	0.305	1.002	0.586	3.688	−1.969
	(1.23)	(2.30)	(2.07)	(1.27)	(0.19)	(0.23)	(0.32)	(63)	(1.17)	(2.13)	(1.01)*	(1.22)
Black share	0.073				0.007				0.009			
	(0.08)				(0.01)				(0.09)			
Black share	−0.001				−0.0001				0.0008			
squared	(0.002)				(0.0003)				(0.002)			
College	−2.171	−1.63	−1.132	−3.652	−0.514	−0.205	−0.531	−1.127	−0.719	−0.515	−1.682	−0.175
	(0.51)*	(0.79)	(0.47)*	(1.01)*	(0.20)*	(0.24)	(0.27)	(0.97)	(0.27)*	(0.29)	(0.59)*	(1.16)
Industry	−0.009	−0.28	0.522	0.817	0.454	0.708	0.114	0.149	0.0831	−0.266	0.354	0.309
	(0.39)	(0.35)	(0.71)	(2.91)	(0.12)*	(0.11)*	(0.19)	(0.32)	(0.22)	(0.27)	(0.38)	(0.49)
Number of states	35	18	13	4	38	21	13	4	38	21	13	4
N	475	149	245	81	1,102	609	377	116	971	478	377	116
Regression sum of squares	23.908	3.36	16.009	13.165	3.335	1.673	1.367	1.526	3.335	9.117	14.454	2.623

Note: Robust standard errors in parentheses. * indicates p-value less than or equal to 5%.

tions. The ratio of the percent of college-educated women to the percent of college-educated white men ($COLL^{FWM}$) is trend stationary—that is, we can reject the null hypothesis of a unit root when a deterministic trend is included. The ratio of blacks in interest rate-sensitive industries (IND^{BWM}) appears to be non-stationary in the basic specification (Column 1), but other tests reveal no evidence of a unit root (Columns 2–4). The only variables with ambiguous test results are the ratio of women's labor force participation rates to white male labor force participation ($LFPR^{FWM}$) and the ratio of white women's labor force participation rates to white male rates ($LFPR^{WFWM}$). The panel is perfectly balanced with regard to these two variables, so we also performed the Im, Persaran, and Shin panel unit root test using various specifications. We rejected the presence of a unit root in all cases. Therefore, we assumed that these two variables were non-stationary for the purposes of our analysis.

The real federal funds rate is the only variable which is invariant across states. Therefore, we use standard augmented Dickey-Fuller (ADF) unit root tests to examine the stationarity of this variable. The Schwartz information criterion was used to determine number of lags. The tests reveal the real federal funds rate to be trend stationary with an ADF test statistic of −5.07 and a p-value of less than 0.002.