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When Work Disappears: Racial Prejudice and Recession Labour Market Penalties

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Abstract
This paper assesses whether racial prejudice and labour market discrimination is counter-cyclical. This may occur if prejudice and discrimination are partly driven by competition over scarce resources, which intensifies during periods of economic downturn. Using British Attitudes Data spanning three decades, we find that prejudice does increase with unemployment rates. We find greater counter-cyclical effects for highly-educated, middle-aged, full-time employed men. For this group, a 1%-point increase in unemployment raises self-reported racial prejudice by 4.1%-points. This result suggests that non-White workers are more likely to encounter racially prejudiced employers and managers in times of higher unemployment. Consistent with the estimated attitude changes, we find using the British Labour Force Survey that racial employment and wage gaps increase with unemployment. The effects for both employment and wages are largest for high-skill Black workers. For example, a 1%-point increase in unemployment increases Black-White employment and wage gaps for the highly educated by 1.3%-points and 2.5%. Together, the attitude and labour market results imply that non-Whites disproportionately suffer during recessions. It follows that recessions exacerbate existing racial inequalities

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"To make matters worse, the current economic and social crises threaten to widen some equality gaps that might have closed in better times." (Equality and Human Rights Commission, 2010)

1. Introduction

Anecdotal evidence suggests that levels of racial prejudice have increased during the recent Great Recession.[1] This is in line with predictions from a theoretical literature that highlights the propensity for prejudice to increase during periods of economic downturn due to increased competition for scarce resources (Levine and Campbell, 1972; Frijters, 1998; Smith, 2012; Caselli and Coleman, 2013). This paper investigates whether self-reported racial prejudice is counter-cyclical. Our measure of racial prejudice is found in British Social Attitudes Surveys between 1983 and 2010, and is a declaration by a respondent of being 'not prejudiced at all', 'a little prejudiced', or 'very prejudiced' against people of other races. To identify the effects of macroeconomic conditions, we exploit variation in unemployment rates across geographic regions and time. Our findings suggest that prejudice does increase with unemployment, with the effect owed to large increases among highly-educated, middle-aged, full-time employed men. For example, it is estimated for this subgroup that a 1%-point increase in the unemployment rate increases self-reported prejudice by approximately 4%-points or 10% relative to mean prejudice levels.

Importantly, the increase in racial prejudice may have deleterious labour market effects for non-White minorities. This follows given that educated middle-aged White men are highly likely to be over-represented amongst employers and managers, as well as having the most political power within firms. Therefore, an increase in racial prejudice among this group may increase labour market discrimination. Increased discrimination could be widespread and affect non-Whites at all levels within firms if there is a general increase in taste-based discrimination, or it could be concentrated among the highly skilled if the propensity to discriminate arises from increased competition among employees with similar positions. We test this hypothesis using data from the 1993-2012 versions of the Quarterly Labour Force Survey (QLFS). To separately identify the effects of racial prejudice from the effects of anti-

[1] Examples from the popular press include: The Telegraph, 19 January 2009; Reuters, September 1 2008; The Times 14 January 2009; The New York Times 12 September 12 2009.

[2] Some work outside economics does examine the determinants of self-reported racial prejudice using the

immigration attitudes, we restrict our analysis to native born individuals. The results suggest that non-Whites are worse off during recessionary periods in terms of employment and earned income. We refer to the increased racial wage gaps during periods of high unemployment as the ‘recession wage penalty’ (RWP), and correspondingly refer to the increased racial employment gaps as the ‘recession employment penalty’ (REP). Interestingly, the RWP is largest for older, high-skilled workers in private firms. Disaggregation by race reveals that Black workers have the largest RWP. For example, the Black-White wage gap for highly educated workers is estimated to increase by 2.5% for every 1%-point increase in unemployment. The REP is also highest for highly skilled workers; however both the Black-White and Asian-White REP are similar. In particular, the results imply that the probability of these groups being employed decreases by 1.3% for every 1%-point increase in the unemployment rate.

For decades economists have developed theories of racial prejudice (Lang and Lehman, 2012; Altonji and Blank, 1999; Arrow, 1998; Becker, 1957) and have empirically examined its economic consequences (Guryan and Kofi Charles, 2013; Lang and Lehmann, 2012; Fryer and Torelli, 2010; Ritter and Taylor, 2011; Lang and Manove, 2011; Dawkins et al, 2005; Lang et al, 2005; Chay, 1998; Card and Krueger, 1992; Donohue and Heckman, 1991). There are also large independent literatures investigating the determinants of social attitudes, including attitudes towards immigrants and immigration policy (Quillian’s, 1995; Dustmann and Preston, 2005; Dustmann and Preston, 2007; Pettigrew 1998; Mayda 2006; Hainmueller and Hiscox 2007; Facchini and Mayda, 2009). However, despite these influential literatures there is little economic research on the determinants of self-reported racial prejudice. Understanding the economic determinants of prejudicial attitudes is therefore academically valuable, as is exploring how racial prejudice may translate into worse labour market outcomes for non-Whites.

Our findings also have implications for policies targeting ethnic minorities residing in Britain. This arises because of already existing inequalities: minorities live in worse housing (Phillips and Harrison; 2010), are taught by lower quality teachers (Clotfelter et al; 2004) and are in worse health (Lordan and Johnston, 2011and Bollini and Harold, 1995). Additionally, the unemployment rate of minorities in Britain has been approximately double that of Whites over the last 40 years, with only half of this gap explained by residential segregation, education differences and other observable factors (Leslie et al, 2001; Blackaby et al, 2002 and Heath and Li, 2007). Higher levels of racial prejudice widen this gap, and thus the recent recession may have reversed some of the gains made during the past decades (Equality and Human Rights Commission, 2010).

The remainder of this paper is organised as follows: In Section 2 we provide a background for our study by discussing theoretical and empirical work related to prejudicial attitudes and discrimination. In Section 3 we describe the British Social Attitudes data, along with our methodology. We also document the results for the empirical work on attitudes. In section 4 we describe the data sources used to consider labour market impacts, methodology and results. The final section is a discussion.

2. Racial Prejudice and Discrimination

For many years economists have been interested in a diverse range of individual attitudes and values (Gentzkow and Shapiro, 2004; Voigtländer and Voth, 2012; O' Rourke and Sinnott, 2001). However, to our knowledge economists have never empirically examined the determinants of self-reported racial prejudice.[2] This is puzzling, because many economic studies examine attitudes towards immigrants and immigration policy (Quillian's, 1995; Dustmann and Preston, 2006; Dustmann and Preston, 2007; Pettigrew 1998; Mayda 2006; Hainmueller and Hiscox 2007; Facchini and Mayda, 2009). Particularly relevant studies include Lahav (2004), who finds that immigration attitudes are related to perceptions of economic conditions, Kesslar and Freeman (2005), who find that as economic conditions worsen so does public opinion towards migrants, and Dustmann and Preston (2007), who demonstrate that racial prejudice and anti-immigration attitudes are strongly related in the UK.

The dearth of economic studies is also puzzling given the importance of prejudicial attitudes in shaping the life chances of ethnic minorities. For example, racial prejudice has been suggested as an important causal factor in determining policies that target minorities (Bobo,

[2] Some work outside economics does examine the determinants of self-reported racial prejudice using the British Social Attitudes Survey. However, this work focuses on the correlations between individual and household characteristics and prejudicial attitudes (Evans 2002; Rothon and Heath 2003; Ford 2008). For example, Ford (2008) reports low self-reported prejudice amongst the highly educated, the professional classes and women.

1991; Sears, 1988). Additionally, lower levels of racial prejudice have been directly linked to support for the Welfare State in Britain (Ford, 2006). Elsewhere, Dustmann and Preston, (2007) find that racial concerns are an important pathway through which opinion towards immigration policies are formed. Prejudice and subsequent discriminatory practices have also been linked to residential segregation (Zubrinsky, 2000 and Zubrinsky and Bobo, 1996), poorer health outcomes (Johnston and Lordan, 2012 and Lauderale, 2006) as well as worse labour market outcomes (Charles and Guryan, 2008 and Goldsmith et al, 2006). Therefore, shifts in racial prejudice during recessionary periods could worsen already existing racial inequalities in a number of domains.

While no economic study empirically examines the determinants of self-reported racial prejudice, there are several related economic literatures. For instance a number of papers examine the macroeconomic determinants of racially-motivated crime. The closest example to our study is Falk et al. (2011), who find a significantly positive relationship between regional unemployment and the incidence of right-wing extremist crime in Germany. The authors hypothesise that the fear of losing a job increases with unemployment, and that this fear lowers tolerance and altruism. Another close example is Antecol and Cobb-Clark (2010), who in contrast find little evidence that racial hostility towards off-base Army personnel is related to local economic vulnerability (measured by the unemployment rate, poverty rate, and income inequality). This finding of weak relationships with macroeconomic conditions is common in the racially-motivated crime literature. For example, macroeconomic conditions are found to be only weakly related to the incidence of anti-foreigner crime in Germany (Krueger and Pischke, 1997), hate crimes in New York (Green et al. 1998), race riots in the US (DiPasquale and Glaeser, 1998), and racial harassment in Britain (Dustmann et al., 2011). We note that physical violence is a very particular and extreme manifestation of racial prejudice, and it is therefore unclear how relevant these results are for our study.

Another related literature theoretically models racial discrimination, scarcity and conflict. For example, Frijters (1998) argues that job uncertainty and scarcity encourages groups of individuals to form coalitions based on observed recognizable characteristics, such as race. Individuals then hire persons from within their own coalition in an attempt to capture all the scarce jobs and ensure future labour market success. Smith (2012) builds on the social identity literature, which finds that people favour members of their own group at the expense of members of other groups, even if securing this outcome creates economic inefficiencies.

He finds that the competition for scarce resources can induce agents without discriminatory attitudes to aggressively discriminate. Finally, Caselli and Coleman (2013) consider ethnic conflict. In their model, ethnicity provides a means for group membership and exclusion, which limits the ability of the losing groups to access the spoils of conflict, such as land titles or government jobs. More generally, there are large literatures in sociology and anthropology based on the theory that discrimination is the result of competition over scarce resources, the effects of which are exacerbated during periods of economic downturn (see Green et al. 1998).

A final related literature considers how the business cycle affects racial wage gaps. Early studies relied on aggregate time series data and examined the ratio of annual earnings for the discriminated group compared to the majority group (Ashenfelter, 1970; Freeman, 1973; O'Neill, 1985). None of these studies found empirical evidence of pro- or counter-cyclical 'pure' wage discrimination. Biddle and Hammermesh (2013) provide the most relevant contribution by considering wage discrimination over the business cycle in the US using the Current Population Survey Merged Outgoing Rotation Groups from 1979 through 2009. They find evidence that there is a counter-cyclical wage gap for African-Americans, but that this is mostly owed to composition effects. They also find an opposite pro-cyclical wage gap for Hispanics, but the effect is only marginally significant. There is also a notable contribution in the immigration literature by Dustmann et al. (2010). The study compares natives and immigrants from OECD and non-OECD countries in Germany and the UK, and highlights counter-cyclical unemployment effects for immigrants that are particularly pronounced for those from non-OECD countries.

Based on this empirical and theoretical research, we hypothesis that macroeconomic conditions can alter an individual's prejudicial attitudes. In particular, we envisage that each individual has a baseline level of prejudice that is affected by the degree to which they are experiencing 'bad economic times' and have to compete over scarce resources. In recessions, when lay-offs and wage cuts are more common, we expect individuals become less tolerant towards people of other ethnicities. This is in line with the Caselli and Coleman (2013) theoretical model, whereby ethnicity provides a means for group membership and exclusion. In addition, it is also an explanation as to why those who are low educated and working have relatively high-level levels of intolerance regardless of the unemployment rate. That is, they are routinely experiencing bad economic times (see Figure 1 which we will subsequently discuss). It is not necessary to make an assumption as to whether long-run baseline prejudice levels are high (becoming low when an individual experiences good economic fortune) or low (becoming high when an individual experiences bad economic fortune). In this work we are simply interested in examining whether prejudice is more prevalent in the community during periods of high unemployment.

We also hypothesise that different sub-groups of society more readily alter their attitudes as unemployment rises. This fits with the aforementioned work and the findings of Mayda (2006), who finds that anti-immigration attitudes heighten when an individual's own circumstances are most threatened by immigration. It also fits with our hypothesis that tolerance is a function of one's own individual circumstances. Thus, those who are experiencing good economic times prior to the recession are likely to become more intolerant towards ethnic minorities, particularly if they are competing alongside these groups in a work environment for income and employment opportunities. Thus, high-skilled individuals would become more prejudiced if ethnic minorities were also high-skilled. In the UK, native ethnic minorities have spent more years overall in school and possess a higher number of degree qualifications than native Whites (Blackaby et al, 2002; Gillborn and Gipps, 1996 and Modood et al, 1997). If we consider the Labour Force Survey data from 1993-2012, which we describe later, Blacks have slightly higher education than Whites in terms of degree attainment, while Asians and other ethnicities have much higher education levels. Additionally, Dustmann and Fabbri (2006) highlight that in the UK immigrants have higher levels of skill and education, than immigrants in other developed countries. Their work also indicates that the immigration attitudes of the educated are most influenced by economic factors.

3. Estimating Racial Prejudice over the Business Cycle

3.1. Data

Our data is drawn from the British Social Attitudes (BSA) Survey, which is a mostly annual series initiated in 1983. The exceptions are 1992 and 1998 when the BSA was not conducted. Each year the BSA Survey asks over 3,000 people questions that gauge their social, political and moral attitudes. Participants are selected using random probability sampling which ensures that the survey is representative of the British population. The key racial prejudice question was included in all years apart from 1993, 1995 and 1997, and states:

"How would you describe yourself? (1) not prejudiced at all, (2) a little prejudiced, or (3) very prejudiced, against people of other races?".

The relative frequencies of responses (1), (2) and (3) are 66%, 30% and 4%, respectively. Given the low frequencies associated with response (3), we aggregate the responses "a little prejudiced" and "very prejudiced" to create a binary outcome variable measuring any prejudice.

Figure 1 presents mean levels of this measure by some key demographics.[3] These graphs demonstrate three interesting features. First, the highly educated report less prejudice. This is consistent with the highly educated being more tolerant (Sullivan and Transue, 1999 and Hello et al, 2006). Second, those not working report less prejudice than those who are working (or equal in the case of highly educated females). Note that the non-working group consists predominantly of retirees and home-makers, with the unemployed a relative minority. Third, males report more prejudice than females. This is consistent with Lubbers et al (2002), and Johnson and Marini (1998). Combined, these relationships imply that low-educated working males are the most prejudiced (42%) and high-educated women are the least prejudiced (25%).

It is likely that self-reported prejudice is a biased measure of true prejudice due to people voicing a different attitude publicly (to the interviewer) than they hold privately. That is, self-reported prejudicial attitudes are likely under-reported (Rudman et al, 2001). However, Evans (2002) highlights that the self-reported prejudice variable in the BSA is correlated closely with other indicators of racial intolerance, and we expect those with the highest level of prejudice to declare. Additionally, the tendency to under-report owing to social desirability is unlikely to be associated with the prevailing unemployment rate, unless the unemployment rate has a strongly significant effect on the level of prejudicial attitudes within the general community. We do note that it is possible that during recessions some media outlets and political groups increasingly discuss ethnic minorities in a critical manner. Subsequently, individuals may be more likely to declare their pre-existing prejudice as they deem such

[3] In Figure 1 those who have achieved a CSE, O level or GCSE are described as having medium education, and those with an A-level or degree are described as highly educated. All others are described as possessing low levels of education.

views as more appropriate. Here we argue that this still represents a change in levels of racial tolerance, given that being willing to declare one's prejudice also implies an increased probability of acting upon it. Finally, the tendency to under-report should bias the estimated relationship between unemployment rates and attitudes downwards (towards zero).

Overall the data comprises 23 survey years, and eleven areas within the UK. These areas are: North East England, North West England, Yorkshire and Humber, East Midlands, West Midlands, South West England, East of England, London, South East England, Wales and Scotland. Official unemployment rate data is only available for these areas from 1992, and so to construct area-level unemployment rates for 1983-1991 we use area-level information on the numbers of claimants for Job Seekers Allowance (JSA). The difficulty is that claimant rates are not necessarily equivalent to unemployment rates, because some of the unemployed are ineligible for or may not claim unemployment benefits. To control for this difference, we construct area-level unemployment rates by multiplying the area-level claimant rate with the ratio of national unemployment rates to national claimant rates. That is, for the period 1983-1991 the unemployment rate used in this work is derived as $cr_{ayt} \cdot ur_{yt}/cr_{yt}$, where cr_{ayt} is the area-level claimant rate, and ur_{ty} and cr_{yt} are the national unemployment and claimant rates, respectively.[4]

Figure 2 presents the national self-reported racial prejudice series across time. It shows a general downward trend across the 1980s and 1990s, followed by an upward trend in the 2000s. The upward trend may have been partly caused by the world-wide increase in terrorism during this period. Another interesting feature is the sharp uptake in the early 1990s. This corresponds with a period of high unemployment in Britain; the national unemployment rate peaked at 10.8% in early 1993. Figure 3 presents the estimated cross-sectional relationship between self-reported racial prejudice and unemployment using between area variation (an Epanechnikov kernel function and a rule-of-thumb bandwidth are used). The non-parametric estimates are presented for four labour market groups: employed, unemployed, retired, and all others, which are primarily home makers. The figure shows a strong positive relationship between racial prejudice and unemployment for workers across all UR values. There is also a positive relationship for the unemployed, but this exists only up

[4] This approach follows advice received from the Office of National Statistics. It assumes that there is no difference between the divergence of claimant and unemployment rates across areas

until an UR of 8%. The relationships are weak for retirees and homemakers. Though it is difficult to draw any firm conclusions from cross-sectional correlations, this figure suggests it may be important to consider heterogeneity in the underlying relationship. This is in line with our hypothesis that self reported prejudice may be more likely to change for some groups in society more than others during recessions.

3.2. Methodology

To model the effects of macroeconomic conditions on the racial prejudice of White respondents, we use a linear regression model with area-specific intercepts and area-specific linear time trends:

$$RP_{iay} = \delta UR_{ay} + X_{iay}\beta + \mu_y + \alpha_a^1 + \alpha_a^2 y + \varepsilon_{iay} \quad (1)$$

where RP_{iay} is the reported racial prejudice of individual i residing in area a in year y, UR_{ay} is the area-level unemployment rate, $X_{iay}$is a vector of individual-level control variables, μ_y is a year fixed-effect, α_a^1 is a area-specific intercept, $\alpha_a^2 y$ is an area-specific time-trend, and ε_{iay} is a random disturbance term. Standard errors are clustered by area to allow for correlation between disturbances across years within areas.

Area-level intercepts (fixed-effects) are included in equation (1) as it is likely that racial prejudice is influenced by local factors. Area-specific linear time trends are also included because it is plausible that there has been a trend across time towards the acceptance of other ethnicities, and that the slope of this trend may differ by area. More generally, area-specific trends can capture time-variant unobservable factors that are associated with both prejudice and unemployment. Given the inclusion of area-level intercepts and area-specific time trends, the effect of the unemployment rate on prejudice is identified by within-area variations in unemployment in relation to within-area variation in prejudice around its trend.[5]

[5] We have also estimated models with area-specific covariates that control for the effects that the September 11 2001 terrorist attacks and subsequent events may have had on racial prejudice. These covariates amount to the addition of the term $\alpha_a^3 \cdot \mathrm{I}(y > 2001)$ to equation (1), where $\mathrm{I}(\cdot)$ is an indicator function that equals one if the argument is true and zero otherwise.

The vector of individual-level control variables (X_{iay}) includes, gender, age, age-squared, number of children, marital status (married, separated/divorced/widowed or single), employment status (full-time employed, part-time employed, unemployed, retired, full time student, with all others (mostly home-makers) as the control group), log of household income and log of household income squared (in 2010 prices). To the extent that voting preference can capture some propensity of an individual to be more or less liberal (Pallage and Zimmermann, 2006), or indeed their personality in general (Schoen and Schumann, 2007 and Caprara et al, 1999), we include dummies indicating if a person either has no party allegiance or votes for: Labour, Alliance, or 'other party' (Conservative is thus the reference group). Appendix Table A1 reports the sample means of these variables by 'not prejudiced' and 'prejudiced'.

Particularly important covariates are the education variables, which are used to define estimation subgroups. Those with A-levels or a university degree are described as highly educated. In the UK, students completing A-levels stay in school until roughly 18 years and generally aim for third level education. Those who have achieved a certificate of secondary education (CSE), O-levels or a general certificate of secondary education (GCSE) are described as having medium education (CSE and O-levels were replaced by the GCSE in 1988). All three qualifications represent a low-level secondary school qualification that is usually achieved when the student is aged 15. Individuals without any qualification are described as possessing low levels of education. This classification approach generates three education groups of roughly equivalent sample size. Importantly, education information is missing from the BSA Survey in years 1983 and 1984, and therefore the regression sample begins in 1985.

3.3. Results

Table 1 presents coefficient estimates for equation (1), separately by males and females. The coefficient on the unemployment rate equals 0.004 for males and is significant at the 10% level. This implies that self-reported prejudice increases by 0.4%-points when the UR increases by 1%-point, suggesting there is only a small increase in male racial prejudice during economic downturns. For females the estimated unemployment effect equals 0.003 and is not statistically significant. We draw similar conclusions if we do not collapse our self-reported prejudice measure (not prejudiced at all, a little prejudiced, very prejudiced) into a binary indicator and instead use an ordered probit model (estimated coefficient for men equals 0.008 and its p-value equals 0.061).

The control variables coefficients in Table 1 highlight that male full-time workers, and female full-time and part-time workers are more likely to be racially prejudiced than the control group.[6] Relative to those with low education, the highly educated are less likely to report racial prejudice, and the medium educated are more likely to report racial prejudice. For income our results imply that middle-income households are most likely to be racially prejudiced. This is evidenced by the inverse-U relationship between income and prejudice. Voting preference is also indicative of prejudicial attitudes. Table 1 reveals that Labour and Alliance voters report being less racially prejudiced than conservative voters regardless of gender. This is also the case for those who report having no political allegiance.[7]

As discussed in Section 2, we are interested in how the prejudicial attitudes of different sub-groups change with the business cycle. In particular, we hypothesise that the counter-cyclical relationship will be strongest for high-skilled workers (e.g. more experienced and better educated). Ethnic minorities who are native to Britain have higher education than White natives, and therefore in periods of economic downturn the labour market competition between ethnic minorities and Whites will be most intense for high-skilled workers. Note that this dynamic is relatively unique to the UK. For example, in the United States both native and immigrant ethnic minorities are lower skilled than White. In this context, studies have concluded that fears about the adverse effects of labour-market competition are the cause of anti-immigration attitudes among low-skilled, blue collar workers (Scheve and Slaughter, 2001; Mayda, 2006).

In Table 2 we present estimated UR coefficients separately by age (18-34, 35-64 and 65+), education level (low, medium, high) and employment status (full-time, full-time or part-time, not employed, retired). Considering age, the only statistically significant effect is for males

[6] This gender difference is perhaps owed to it being more common in the UK for females to be part-time employed than in other countries and in comparison to males (Manning and Petrongolo, 2008)
[7] It is possible that political allegiance may be driven by prejudicial attitudes and as such this set of variables may be endogenous. Omitting them from the regression models, however, has very little impact on the remaining estimated coefficients.

aged 35-64. For this group, the estimate of 0.017 implies racial prejudice increases by 1.7%-points for every 1%-point increase in UR. Or to put it another way, for a 4%-point increase in UR (as seen in the most recent recessionary period) it is estimated that 7 additional men out of every 100 self-report racial prejudice. The sub-analysis by education reveals that racial prejudice is indeed most strongly counter-cyclical for the high education group. The estimated effects for males and females equal 0.013 and 0.016. With respect to employment status, Table 2 reveals that prejudice attitudes are counter-cyclical for working males. If we consider full-time working males, the results imply that a 4%-point increase in UR increases racial prejudice by almost 7%-points.[8]

To explore this finding further, the bottom panel of Table 2 presents additional sub-group analysis based on interactions between age, education and employment. The results show that for men the UR effect is large and statistically significant for each interacted sub-group. The effect is particular large in the final row, which presents estimates for the highly educated that are full-time employed and aged 35-64. The estimate suggests that a 4%-point increase in UR increases racial prejudice by 16%-points. This is a 45% increase relative to the mean racial prejudice for this group of 36%. For females, racial prejudice is also most strongly counter-cyclical for the highly educated that are full-time employed and aged 35-64. The estimate suggests that a 4%-point increase in UR increases racial prejudice by 8%-points. Given that the average individual in this sub-group is relatively high-skilled, these finding concur with our hypothesis that racial prejudice is most strongly counter-cyclical for high-skilled workers.

Importantly, an alternative explanation is possible for the results presented in Tables 1 and 2. Prejudice levels amongst workers may increase during recessions if individuals with relatively low levels of racial prejudice are more likely to become unemployed than individuals with high levels of racial prejudice. This systematic selection out of employment is plausible if prejudiced workers cooperate together and compete against more tolerant individuals for scarce resources. In this scenario, individual-level prejudice may not have increased during recessions; however, it is still the case that the likelihood of encountering a

[8] Including area-specific controls for the sharp increase in racial prejudice post September 11 2001 does not alter these results.

prejudiced colleague, manager or employer has increased. Thus, we still view this as a worrisome change.

In Tables 1 and 2 the unemployment rate is defined as the average rate of unemployment for all individuals residing in a particular area in a particular year. This may not, however, be the most suitable measure for each subgroup. It is possible that the relevant measure for a subgroup is not the area-year average across all individuals, but rather the area-year average across only those individuals belonging to the subgroup (e.g. unemployment rate for highly educated men aged 35-64). This measure may more adequately capture labour market competition. Unfortunately, heterogeneous unemployment rates by region-year are unavailable; the areas and survey years in the British Labour Force Survey only partially match those in the BSA, preventing us from performing our own calculations. Nevertheless, we can demonstrate that the area-year average UR is a significant predictor of employment outcomes for all subgroups and is therefore relevant. For example, using the same fixed-effect methodology as used in Tables 1 and 2, we find that the unemployment rate has a significantly negative effect on the full-time employment probabilities of low (-0.029), medium (-0.011) and high (-0.006) education subgroups.

The results in Tables 1 and 2 are also robust to alternative model specifications (see Appendix Table A2). First, the continuous measure of UR was replaced with dummies representing the UR groupings of 5-8%, 8-11% and >11% (<5% was the omitted comparison group). The estimated effects for the high-skilled subgroup (i.e. highly educated, full-time employed and aged 35-64) equal 0.032, 0.087 and 0.183. Second, lagged UR variables (UR_{at-1}, UR_{at-2}) were added to the specification. The dynamic specifications suggest that the UR effect is largely contemporaneous. For the high-skilled subgroup the estimated effects for UR_{at}, UR_{at-1} and UR_{at-2} equal 0.041, -0.003 and 0.007, respectively. Finally, using a three year average of UR, which should reduce measurement error, counter-intuitively reduces the estimated UR effect for some sub-groups; however, the effect for the high-skilled subgroup remains strong (0.046).

Finally, we can show that unemployment also has an effect on expectations of future prejudice rates in Britain. In addition to the question on own racial prejudice, the BSA asks respondents “Do you think there is generally more racial prejudice in Britain now than there was 5 years ago?” with the respondent given the following response options: “less”, “more” or "about the same amount"?, and "Do you think there will be more, less, or about the same amount of racial prejudice in Britain in 5 years time compared with now?". In Appendix Table A3 we present the estimated effects of UR on a binary variable indicating that the respondent believes there is less prejudice now than 5 years ago, and on a binary variable indicating that the respondent believes there will be less prejudice in 5 years time. The results show that an increase in UR significantly reduces the probability that the respondent believes there will be less prejudice in 5 years time: the estimated effect for the high-skilled subgroup defined above equals -0.017 (relative to a mean of 0.22). In other words, higher unemployment rates now lead people to expect higher levels of racial prejudice in the future.

4. Estimating Labour Market Discrimination over the Business Cycle

The results in the previous section highlight that males aged 35-64 that are highly educated workers have the most pronounced counter-cyclical racial prejudice. Or, at least there is a larger proportion of intolerant highly educated males in employment during recessions. Assuming these individuals are more likely to be managers, bosses and/or have political power within organisations, should they act on their increased taste for discrimination, this will translate into worse labour market outcomes for non-Whites during periods of high unemployment. In other words, counter-cyclical racial prejudice could lead to counter-cyclical labour market discrimination. The worsening may perpetrate all levels of the organisation if there is an increase in 'pure' discrimination, however it will be concentrated among the highly skilled if the propensity to discriminate arises from increased competition. That is, in the first case those with increased prejudice discriminate against anyone who is of a particular ethnic minority regardless of personal gains, as they gain utility from their discriminatory actions. In the second case, discrimination is targeted towards those at their own skill level. This may be viewed as 'rational' discrimination, because if successful it results in the person's own job and the jobs of their group members being more secure (Frijters, 1998).

It is important to note that the racial labour market gaps that we can identify here could be counter-cyclical in the absence of counter-cyclical racial prejudice. An alternative explanation is that the costs of discriminatory behaviour could be lower during periods of high unemployment (Biddle and Hamermesh, 2013; Baert et al. 2013). Hiring a less qualified White applicant generates costs in terms of foregone production, and turning away a qualified non-White applicant generates additional search costs. During economic downturns there is a greater pool of White applicants to employ, and therefore employers with a taste for discrimination can more easily find qualified White workers.

4.1. Data

Our data source for measuring labour market discrimination is the 1993-2012 versions of the Quarterly Labour Force Survey (QLFS). The QLFS is the main survey of individual economic activity in the UK, and provides the official measure of the national unemployment rate. We consider wage and employment differences between native Whites and native non-Whites, and exclude all immigrants from the analysis. We make this selection decision because non-White immigrants have significantly lower observed and unobserved human capital than native Whites and non-Whites; they generally have lower levels of formal education, lack English language skills, are unfamiliar with local customs and work in unskilled jobs (Berthoud, 2000; Heath and Cheung, 2007). In addition, the composition of the immigrant work force is likely to change over the business cycle, with average immigrant human capital levels moving with the needs of the economy. This arises given that natives are likely to have more ties to their area of residence than non-natives. Individuals are identified as being native to the UK based on their country of birth and we include all respondents aged 18-65 years. We choose to focus on both employment and income as it is plausible that there is heterogeneity in the results for these outcomes across worker groups. For example, the recession wage penalty (RWP) may be worse for highly educated non-Whites whilst the recession employment penalty (REP) may be worse for non-Whites with low education levels. For the purpose of our analysis we generate quarterly unemployment rates specific to 19 geographic regions.[9]

4.2. Methodology

To model the effects of macroeconomic conditions on ethnic employment and wage differentials, we use a linear regression model with intercepts specific to each region of

[9] The geographic regions are: Tyne and Wear; Rest of North East, Greater Manchester, Merseyside, Rest of North West, South Yorkshire, West Yorkshire, Rest of Yorkshire and Humberside, East Midlands, West Midlands Metropolitan County, Rest of West Midlands, East of England, Inner London, Outer London, South East, South West, Wales, Strathclyde, and Rest of Scotland.

residence and quarter:

$$Ln(y_{iaq}) = \delta NW_{ijq} + \gamma(UR_{aq} \cdot NW_{iaq}) + X_{iaq}\beta + \mu_{aq} + \varepsilon_{iaq} \qquad (2)$$

where y_{iaq} is either the wages of individual i residing in area a in quarter q or an indicator as to whether they are in employment, NW_{ijq} is a dummy representing non-White workers (disaggregated racial identifiers are used in some specifications), X_{iaq} is a vector of individual-level control variables, μ_{aq} is an area-quarter fixed-effect, and ε_{iaq} is a random disturbance term. Standard errors are clustered by area-quarter to allow for correlation between individuals within the same area and quarter.

Area-quarter intercepts (fixed-effects) are included in equation (2) to control for any differences in wages and employment across areas and time – this approach amounts to the inclusion of 1482 area-quarter dummy variables. This approach is feasible because the main variable of interest is the interaction term $UR_{aq} \cdot NW_{iaq}$, which varies across individuals within the same area-quarter. The large number of area-quarter terms does however rule out the use of a probit modelling approach for the employment outcome. Consequently, we use linear models for both employment and log wage outcomes. Note also that the pure unemployment rate effect on wages is not identified, as all individuals in the same area-quarter face the same rate. We feel it is preferable to adequately control for differences in outcomes across areas and time, than to include a higher-level fixed-effect term (such as area level intercepts) and identify the pure unemployment rate effect.[10]

A coefficient of interest in equation (2) is δ, which provides a measure of the wage gap between Whites and non-Whites. Typically, the racial wage gap is negative, indicating that non-Whites earn less than Whites; though, given we have omitted all immigrants this may not hold true in our analysis. Importantly, the interaction term is constructed such that δ represents the wage gap at mean levels of UR. Another coefficient of interest is γ, which we've labelled the RWP in the case of wages and the REP in the case of employment. In the case of wages, this term measures how the percentage difference in wages between native

[10] We note that all conclusions are robust to relaxing this assumption, with the estimates from this preferred model being the most conservative.

Whites and non-Whites (wage gap) changes with each 1%-point change in UR. A negative value of γ means an increase in UR worsens the wage gap between Whites and non-Whites, implying counter-cyclical discrimination. Conversely, a positive value of γ implies pro-cyclical discrimination. In a similar vein, the REP measures how the percentage difference in the probability of being employed between native Whites and non-Whites (employment gap) changes with each 1%-point change in UR.

The vector of individual-level control variables (X_{iat}) includes, age, age squared, education level (low, (control group), medium and high), marital status (single (control group), married, separated/divorced/widowed) and the number of children. For the wage regressions only we also include usual hours and usual hours squared as controls. These hours covariates are included because there is a tendency for ethnic minorities to participate in part-time work more frequently than Whites (Blackaby et al, 2002); though in practice they make little quantitative difference to the estimates of γ.

4.3. Results

Table 3 presents estimates of equation (2) for the employment and log wage outcomes, separately by males and females.[11] The four coefficients pertaining to the non-White indicator suggest that non-Whites have worse labour market outcomes than Whites overall. Specifically, non-White males are 14%-points less likely to be employed and their wages are 11% lower than White males (at mean levels of UR). Non-White females are 9%-points less likely to be employed and their wages are 2.5% lower than White females. However, these estimated gaps must be interpreted with caution as they may not solely be caused by discrimination (taste-based or statistical), even when considering a sample of natives only. Alternative explanations are that non-Whites have lower levels of unobserved human capital and that non-Whites systematically prefer occupations or sectors with lower average pay (due to differences in tastes or risk preferences).

Coefficients on the non-White UR interaction term are also significantly negative. For males the employment gap (14%-points) and wage gap (11%) increase by 0.6%-points and 1.2% for

[11] The estimated wage effects of our control variables are all consistent with previous studies. Age has a negative quadratic correlation, and education, marriage and children have a positive correlation.

a 1%-point increase in UR. For females, the employment gap (9%-points) is not affected by UR, while the wage gap (2.5%) increases by 1.4% for a 1%-point increase in UR.[12] For a 4%-point increase in UR (as seen in the recent recession) these estimates imply that relative wages decrease by about 5% for both genders. These estimates are less likely to suffer from omitted variable bias (composition effects) than the estimated average gaps, because unobserved human capital or preference differences between native Whites and non-Whites are unlikely to be associated with differences in area-level unemployment rates.[13] This contrasts with analogous immigration studies in which the systematic self-selection of immigrants into high wage areas is a genuine concern. An approach used in the immigration literature to mitigate the bias is to examine only the sub-group of immigrants who have been living in the country for many years (e.g. Dustmann et al., 2010). This approach is based on the assumption that the current residential location of these established immigrants is less likely to be based upon cyclical macroeconomic conditions. This is an assumption similar to our own. Another approach is to examine more homogenous groups of employees. The idea is that the likelihood of omitted variable bias (unobserved human capital differences) is lessened if identification comes from comparisons between workers in similar age, education and occupation groups. We implement this approach below.

Table 4 presents the results from a sub-group analysis for men (age, education, occupation, work hours, sector). We focus on males given that the employment selection issues are far less severe than for females. The results between genders are however generally consistent. The equivalent female results are included in Appendix Tables A4 and A5. In Table 4 we report the coefficient on the non-White dummy variable that captures the racial wage/employment gap and the interaction term that captures the RWP/REP. The non-White dummy variable is negative and significant for all sub-groups in both the employment and wage models. For employment it is particularly large for the 18-34 age group (-0.156) and for the medium education group (-0.176). For wages it is particularly large for the 35-64 age group (-0.124), medium and high education groups (-0.105 and -0.104) and for workers in private firms (-0.120).

[12] The results in Table 4 pertain to the probability of being employed, but for some readers it may be also interesting to consider the association with unemployment. Although not shown here, the estimates are very similar in magnitude (but opposite in sign) to the employment results.

[13] More specifically, our estimates would be partly driven by bias if non-White workers with high unobserved human capital systematically move to relatively low UR areas at higher rates than White workers and non-White workers with low unobserved human capital.

More relevant are the estimated coefficients on the interaction term UR · Non-White. For both employment and wages, this interaction term is larger for older individuals (-0.008 and -0.017) and highly educated individuals (-0.012 and -0.015). The estimated wage effects (RWPs) are also larger for high-skilled non-manual workers (-0.014) and private workers (-0.016). The difference between part-time and full-time workers is small and not statistically significant. Overall, these results suggest that high-skilled non-White workers, whether defined by experience (age), education or occupation, suffer the most during recessions.[14] If we estimate the RWP for the combined high-skill group – aged 35-64, high education levels, high-skill non-manual occupation – the estimated RWP equals 2.7% (p-value = 0.003). This estimate implies that male non-White relative wages decrease by around 11% with a 4%-point increase in UR, and increase the racial wage gap from a comparatively low 6% to a high 17%.

Next we explore the potential for heterogeneity across the non-White sub-groups by disaggregating this broad classification into three groupings: Blacks, Asians (Indian, Pakistani or Bangladeshi) and other ethnicities (including Chinese).[15] Practically, we replace the non-White dummy variable in equation (2) with dummy variables representing these three sub-groups. Similarly, we replace the interaction between non-Whites and UR with three interaction terms. The results are documented in Table 5 for our full sample as well as a sub-analysis for older and highly educated individuals. These two subgroups were chosen because their estimated effects in Table 4 were particularly large.

The disaggregated results reveal that the effect of the unemployment rate (interaction terms) on employment and log wages is larger for Black men (-0.011 and -0.014) than for Asian men (-0.004 and -0.010) and other non-White men (-0.003 and -0.011). The difference between the non-White sub-groups is particularly large for highly educated workers. The coefficient of -0.025 for the UR · Black interaction term implies that the Black relative wages decrease by 2.5% for a 1%-point increase in the unemployment rate. For a 4%-point increase

[14] Using the natural log of the unemployment rather than UR, as is familiar in the wage curve literature, does not alter these results.
[15] We follow the common practice used in British surveys of labelling individuals from the Indian subcontinent countries of India, Pakistan and Bangladesh as ‘Asian’, and treating these individuals separately to those from other countries in Asia (e.g. China).

in UR (as seen in the most recent recessionary period) it is estimated that the Black wage gap increases from 19% (at average UR levels of 6.4) to 29%. These wage effects are much larger than for Asian men (-0.010) and other non-White men (-0.005).

Conclusions

This paper investigates whether self-reported racial prejudice is counter-cyclical. We find only small general population increases in prejudice during periods of high unemployment. However, we find large increases for White males who are full-time employed, have high levels of education and are middle aged. The increase is especially large for the interaction of these groups. The effect for females who are middle aged, highly educated and in full-time employment is also significant, albeit half the size of the male effect. If these individuals are employers, managers, or have political power within organisations, and act on their increased taste for discrimination, this will translate into worse labour market outcomes for non-Whites during periods of high unemployment. We note that these results may be also be explained by more tolerant high-skill individuals losing their employment during recessions; however, even in this case, the results imply that there are relatively more prejudiced workers during recessions. This attitude shift may affect ethnic minorities at all levels of an organisation if there is an increase in 'pure' discrimination, however it will be concentrated among the highly skilled if the propensity to discriminate arises from increased competition only. Therefore, the second part of the paper investigates whether racial labour market gaps are significantly related to unemployment rates, and how the relationship varies across subgroups.

Our empirical analysis of the Quarterly Labour Force Survey (QLSF) highlights that there are significant differences in employment and wages between Whites and non-Whites, and that these gaps significantly increase with the unemployment rate (counter-cyclical). Specifically, we find the recession employment and wage penalties are the greatest for non-White workers with high skill levels, particularly for males. For example, the estimated RWP for men aged 35-64 with high education levels working in high-skill non-manual occupations equals 2.7%, implying that non-White relative wages decrease by around 11% with a 4%-point increase in UR, as seen in the recent Great Recession. Further analysis reveals that Black workers suffer larger RWPs than other non-White workers.

Overall, the attitudes and labour market results support the hypothesis that racial prejudice is driven partly by increased labour market competition between racial groups during periods of high job uncertainty and scarcity.[16] That is, the comparatively large effects for high-skill subgroups with respect to both self-reported racial prejudice and racial labour market gaps suggest that high-skill Whites may be utilising racial discrimination as a means to retain employment and achieve high wages.

Importantly, we are unable to exclude two additional explanations for the QLFS results. First, the costs of discriminatory behaviour could be lower during periods of high unemployment due to the greater pool of qualified White applicants to employ and promote. Second, the QLFS results could be driven by a converse process, in which affirmative action policies are promoted during economic expansions and abandoned during periods of economic recession. Both of these explanations imply that racial labour market gaps are counter-cyclical even in the absence of increased racial prejudice (tastes for discrimination). Though possible, in light of our self-reported racial prejudice results, we believe these additional processes could only be partially driving the RLMPs.

Regardless of the exact underlying processes, we robustly find that during recessions White workers are more likely to be racially prejudiced and existing racial inequalities in the labour market widen. The recent commentary in the popular press and the corresponding statements from equality and human rights groups appear correct. This is despite native non-Whites achieving similar education levels to Whites in the UK. Given that non-Whites continue to experience significant inequalities in health, housing and schooling quality, we argue that policy makers must be mindful of how recessions can disproportionately penalise minority individuals and should develop policies to avoid these harmful effects in the future.

[16] These conclusions are robust to the consideration of a number of non-linear specifications, however this dynamic may be specific to the UK where ethnic minorities receive a similar number of third level qualifications in comparison to Whites.

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Figure 1: Self-Reported Racial Prejudice by Education, Gender and Employment Status

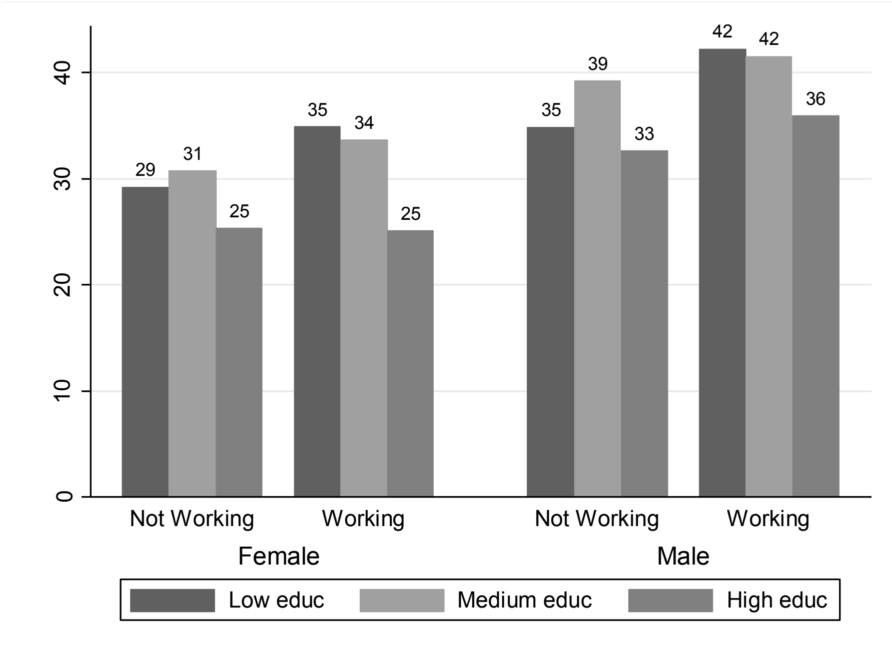

Figure 2: Self-Reported Racial Prejudice across Time by Gender

Figure 3: Cross-Sectional Relationship between Regional Unemployment Rates and Self-Reported Racial Prejudice

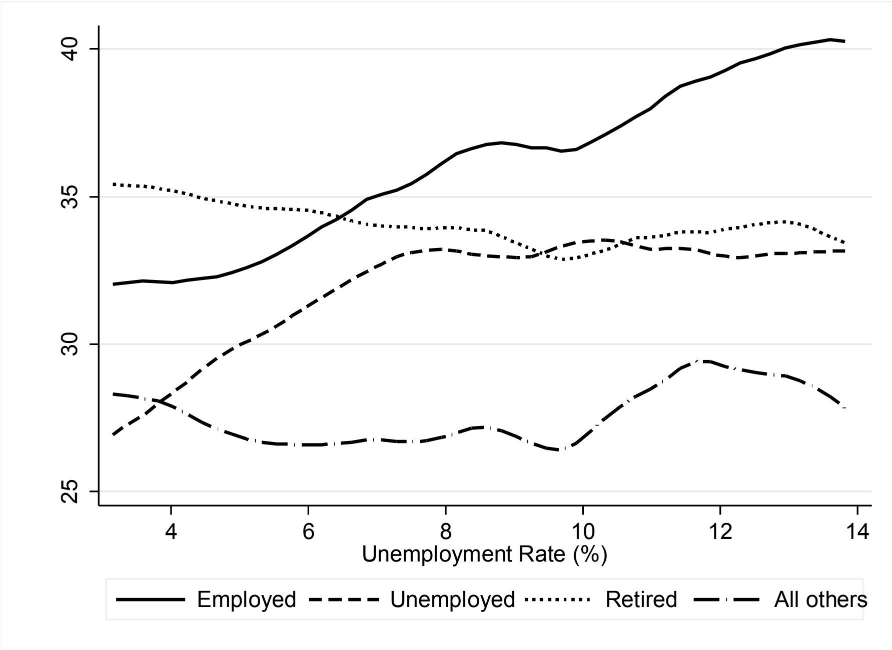

Table 1: Probit Models of Self-Reported Racial Prejudice by Gender

	Males		Females	
	ME	SE	ME	SE
Unemployment rate	0.004*	(0.002)	0.003	(0.005)
Age	-0.003**	(0.001)	0.001	(0.001)
Age squared / 100	0.003**	(0.001)	-0.001	(0.001)
Number of children	-0.011***	(0.004)	-0.015***	(0.004)
Married	0.065***	(0.011)	0.006	(0.011)
Separated or divorced	0.028***	(0.010)	-0.001	(0.015)
Widowed	0.002	(0.023)	-0.029**	(0.012)
Full-time employment	0.038**	(0.017)	0.025**	(0.012)
Part-time employment	0.010	(0.018)	0.029***	(0.010)
Unemployed	0.028	(0.018)	0.025*	(0.013)
Retired	0.022	(0.021)	0.032***	(0.009)
Full-time student	-0.054*	(0.031)	-0.078***	(0.014)
Education medium	0.024*	(0.013)	0.017*	(0.010)
Education high	-0.034**	(0.017)	-0.059***	(0.013)
Log income	0.219***	(0.031)	0.104***	(0.034)
Log income squared	-0.043***	(0.007)	-0.016**	(0.007)
No party allegiance	-0.032***	(0.010)	-0.044***	(0.009)
Labour voter	-0.108***	(0.008)	-0.097***	(0.010)
Alliance voter	-0.110***	(0.011)	-0.081***	(0.011)
Other party voter	0.074	(0.055)	-0.004	(0.040)
Mean outcome	0.375		0.294	
Pseudo R-squared	0.029		0.031	
Sample size	17641		22066	

Note: Figures are probit marginal effect estimates (ME). Dependent variable equals one if a little prejudiced or very prejudiced, and zero otherwise. Included in the model but not shown are area-level fixed-effects, area-level linear time trends, and year fixed-effects. Standard errors clustered at the area level are shown in parentheses. *, ** and *** denote statistical significance at 0.10, 0.05 and 0.01 levels, respectively.

Table 2: Probit Marginal Effects of Unemployment Rate by Subgroups

	Males		Females	
	ME	SE	ME	SE
Age				
(1) 18-34	-0.008	(0.008)	0.006	(0.009)
(2) 35-64	0.017***	(0.004)	0.003	(0.007)
(3) 65+	-0.018	(0.011)	0.000	(0.006)
Education				
(4) Low (no qualifications)	-0.011	(0.011)	-0.007	(0.005)
(5) Medium (CSE / o-levels)	0.011	(0.009)	0.004	(0.008)
(6) High (a-levels / degree)	0.013**	(0.005)	0.016*	(0.009)
Employment				
(7) Full-time	0.017***	(0.006)	0.010	(0.008)
(8) Full-time or part-time	0.013**	(0.006)	0.009	(0.007)
(9) Not employed	-0.010	(0.008)	-0.002	(0.008)
(10) Retired	-0.013	(0.009)	0.010	(0.010)
Interactions				
(11) 35-64 + high educ	0.032***	(0.011)	0.014	(0.011)
(12) 35-64 + full-time emp	0.027***	(0.006)	0.012*	(0.007)
(13) high educ + full-time emp	0.020***	(0.006)	0.013	(0.008)
(14) 35-64 + high educ + full-time emp	0.041***	(0.014)	0.020**	(0.009)

Note: Figures are probit marginal effect estimates for the unemployment rate. Dependent variable equals one if a little prejudiced or very prejudiced, and zero otherwise. Included in all models are the covariates shown in Table 1, area-level fixed-effects, area-level linear time trends, and year fixed-effects. Standard errors clustered at the area level are shown in parentheses. *, ** and *** denote statistical significance at 0.10, 0.05 and 0.01 levels, respectively.

Table 3: Employment and Log Wage Regression Models by Gender

	Males		Females	
	Employed	Log Wage	Employed	Log Wage
Non-White	-0.135***	-0.110***	-0.091***	-0.025***
	(0.002)	(0.006)	(0.003)	(0.005)
UR · Non-White	-0.006***	-0.012***	0.001	-0.014***
	(0.001)	(0.002)	(0.001)	(0.002)
Age	0.052***	0.080***	0.047***	0.053***
	(0.000)	(0.001)	(0.001)	(0.000)
Age squared	-0.001***	-0.001***	-0.001***	-0.001***
	(0.000)	(0.000)	(0.000)	(0.000)
Education medium	0.117***	0.235***	0.150***	0.212***
	(0.001)	(0.002)	(0.002)	(0.002)
Education high	0.166***	0.608***	0.238***	0.647***
	(0.002)	(0.002)	(0.002)	(0.002)
Married	-0.086***	0.192***	-0.299***	0.233***
	(0.007)	(0.013)	(0.007)	(0.012)
Separated / divorced	0.153***	0.142***	0.054***	0.033***
	(0.001)	(0.002)	(0.001)	(0.002)
Number of children	0.010***	0.065***	0.006***	-0.006***
	(0.002)	(0.003)	(0.001)	(0.002)
Sample size	2234250	435804	2633253	463550

Note: Figures are estimated coefficients from linear regression models with 1482 time-region fixed-effects. Log wage models additionally control for work hours and work hours squared. Standard errors clustered at the time-region level are shown in parentheses. *, ** and *** denote statistical significance at 0.10, 0.05 and 0.01 levels, respectively.

Table 4: Employment and Log Wage Models for Males by Subgroups

	Employed				Log wage			
	Non-White		UR · Non-White		Non-White		UR · Non-White	
Age								
(1) 18-34	-0.156***	(0.003)	-0.003***	(0.001)	-0.104***	(0.007)	-0.011***	(0.003)
(2) 35-64	-0.069***	(0.004)	-0.008***	(0.001)	-0.124***	(0.010)	-0.017***	(0.004)
Education								
(3) Low	-0.126***	(0.006)	-0.004**	(0.002)	-0.077***	(0.014)	-0.004	(0.005)
(4) Medium	-0.176***	(0.003)	-0.001	(0.001)	-0.105***	(0.007)	-0.011***	(0.003)
(5) High	-0.069***	(0.003)	-0.012***	(0.001)	-0.104***	(0.010)	-0.015***	(0.004)
Occupation								
(6) All manual	-		-		-0.092***	(0.008)	-0.005	(0.003)
(7) LS non-manual	-		-		-0.062***	(0.011)	-0.011**	(0.005)
(8) HS non-manual	-		-		-0.085***	(0.010)	-0.014***	(0.004)
Work hours								
(9) Part-time	-		-		-0.081***	(0.016)	-0.013**	(0.006)
(10) Full-time	-		-		-0.106***	(0.006)	-0.011***	(0.003)
Sector								
(11) Private	-		-		-0.120***	(0.007)	-0.016***	(0.003)
(12) Public	-		-		-0.076***	(0.011)	-0.011**	(0.004)

Note: Figures are estimated coefficients on a non-White dummy variable and on the interaction term UR·Non-White from linear regression models with 1482 time-region fixed-effects. Rows (7) and (8) present results for 'low skill' and 'high skill' non-manual workers, respectively. Also included but not shown are covariates representing age, educational attainment, marital status and children. Standard errors clustered at the time-region level are shown in parentheses. *, ** and *** denote statistical significance at 0.10, 0.05 and 0.01 levels, respectively.

Table 5: Disaggregated Racial Wage Gaps for Males

	Employed			Log Wage		
	All	Aged 35-64	High Education	All	Aged 35-64	High Education
Black	-0.111***	-0.058***	-0.079***	-0.131***	-0.141***	-0.187***
	(0.004)	(0.005)	(0.006)	(0.009)	(0.012)	(0.018)
Asian	-0.150***	-0.088***	-0.063***	-0.102***	-0.121***	-0.068***
	(0.003)	(0.006)	(0.004)	(0.008)	(0.018)	(0.012)
Other ethnicity	-0.133***	-0.063***	-0.071***	-0.088***	-0.074***	-0.094***
	(0.006)	(0.009)	(0.008)	(0.014)	(0.025)	(0.020)
UR · Black	-0.011***	-0.009***	-0.013***	-0.014***	-0.019***	-0.025***
	(0.001)	(0.002)	(0.003)	(0.003)	(0.005)	(0.006)
UR · Asian	-0.004***	-0.010***	-0.012***	-0.010***	-0.019**	-0.010*
	(0.001)	(0.002)	(0.002)	(0.004)	(0.008)	(0.005)
UR · Other	-0.003*	-0.001	-0.009***	-0.011*	-0.011	-0.005
	(0.002)	(0.003)	(0.002)	(0.006)	(0.010)	(0.009)
Sample size	2234250	1441422	557760	435804	284075	131904

Note: Figures are estimated coefficients from linear regression models with 1482 time-region fixed-effects. Also included but not shown are covariates representing age, educational attainment, marital status and children. Standard errors clustered at the time-region level are shown in parentheses. *, ** and *** denote statistical significance at 0.10, 0.05 and 0.01 levels, respectively.

Appendix Table A1: Description of Covariates Used in Analyses of Self-Reported Racial Prejudice

Variable	Full Sample	Not Prejudiced	Little / very Prejudiced
Male	0.444	0.414	0.505
Age	48.57	48.31	49.11
Age squared / 100	26.86	26.63	27.33
Number of children	0.535	0.550	0.505
Married	0.598	0.579	0.636
Separated or divorced	0.115	0.120	0.104
Widowed	0.114	0.119	0.103
Full-time employment	0.424	0.407	0.457
Part-time employment	0.106	0.109	0.100
Unemployed	0.048	0.049	0.046
Retired	0.223	0.219	0.232
Full-time student	0.019	0.022	0.012
Education medium	0.280	0.267	0.305
Education high	0.369	0.383	0.339
Log income	2.578	2.559	2.615
Log income squared	7.177	7.109	7.314
No party allegiance	0.107	0.107	0.106
Labour voter	0.326	0.351	0.273
Alliance voter	0.115	0.122	0.102
Other party voter	0.026	0.025	0.029
Sample size	39707	26615	13092

Note: All figures are sample means. The omitted (baseline) dummy variables are: never married; looking after the home; no educational qualification; Conservative party preference.

Appendix Table A2: Non-Linear and Dynamic Specifications for Middle Aged, Highly Educated, Full-Time Employed Samples

	All Females (1)	All Males (2)	Males Aged 35-64 (3)	Males with High Education (4)	Males Full-time Employed (5)	Interaction of Groups (3)-(5) (6)
(A) Nonlinear						
$5 < UR_t \leq 8$	0.004	0.020*	0.026	0.008	0.055***	0.032
	(0.016)	(0.011)	(0.018)	(0.015)	(0.012)	(0.039)
$8 < UR_t \leq 11$	0.007	0.017	0.035	0.004	0.077**	0.087*
	(0.024)	(0.024)	(0.039)	(0.018)	(0.030)	(0.054)
$11 < UR_t$	0.052*	0.027	0.057	0.050*	0.105**	0.183**
	(0.027)	(0.030)	(0.053)	(0.027)	(0.048)	(0.072)
(B) Dynamic						
UR_t	0.002	0.003	0.011*	0.019***	0.018**	0.041**
	(0.007)	(0.004)	(0.006)	(0.007)	(0.008)	(0.019)
UR_{t-1}	0.003	0.003	0.012**	-0.012	-0.001	-0.003
	(0.010)	(0.005)	(0.006)	(0.009)	(0.007)	(0.018)
UR_{t-2}	-0.004	-0.006	-0.005	0.000	-0.003	0.007
	(0.008)	(0.008)	(0.009)	(0.016)	(0.008)	(0.017)
(C) Averaged						
$\left(\Sigma_{k=0}^{2} UR_{t-k}\right)/3$	0.002	0.002	0.022***	0.007	0.016	0.046***
	(0.008)	(0.006)	(0.004)	(0.010)	(0.011)	(0.013)
Sample size	22066	17641	9245	7346	10433	3352

Note: Figures are probit marginal effect estimates for the unemployment rate. Dependent variable equals one if a little prejudiced or very prejudiced, and zero otherwise. Included in all models are the covariates shown in Table X, area-level fixed-effects, area-level linear time trends, and year fixed-effects. Standard errors clustered at the area level are shown in parentheses. *, ** and *** denote statistical significance at 0.10, 0.05 and 0.01 levels, respectively.

Appendix Table A3: Probit Marginal Effects of Unemployment Rate on Relative Prejudice Rates

	All Females (1)	All Males (2)	Males Aged 35-64 (3)	Males with High Education (4)	Males Full-time Employed (5)	Interaction of Groups (3)-(5) (6)
(A) Less racial prejudice now compared to 5 years ago						
	0.001	-0.004	-0.016**	-0.001	-0.009	-0.015
	(0.006)	(0.004)	(0.008)	(0.009)	(0.006)	(0.012)
(B) Less racial prejudice in 5 years compared to now						
	0.002	-0.006**	-0.016***	-0.004	-0.013**	-0.017***
	(0.005)	(0.003)	(0.004)	(0.008)	(0.006)	(0.005)
Sample size	20986	16982	8951	7090	10137	3267

Note: Figures are probit marginal effect estimates for the unemployment rate. Dependent variable equals one if the respondent agrees with the statement and zero otherwise. Included in all models are the covariates shown in Table 1, area-level fixed-effects, area-level linear time trends, and year fixed-effects. Standard errors clustered at the area level are shown in parentheses. *, ** and *** denote statistical significance at 0.10, 0.05 and 0.01 levels, respectively.

Appendix Table A4: Employment and Log Wage Models for Females by Subgroups

	Employed				Log wage			
	Non-White		UR · Non-White		Non-White		UR · Non-White	
Age								
(1) 18-34	-0.098***	(0.003)	0.004***	(0.001)	-0.023***	(0.006)	-0.012***	(0.002)
(2) 35-64	-0.050***	(0.004)	-0.002	(0.002)	-0.021**	(0.009)	-0.016***	(0.004)
Education								
(3) Low	-0.111***	(0.006)	0.002	(0.002)	0.039***	(0.015)	-0.003	(0.006)
(4) Medium	-0.115***	(0.004)	0.005***	(0.001)	0.011	(0.007)	-0.013***	(0.003)
(5) High	-0.058***	(0.003)	-0.004***	(0.001)	-0.059***	(0.008)	-0.019***	(0.003)
Occupation								
(6) All manual	-		-		0.003	(0.011)	-0.015***	(0.005)
(7) LS non-manual	-		-		0.009	(0.007)	-0.009***	(0.003)
(8) HS non-manual	-		-		-0.046***	(0.008)	-0.013***	(0.003)
Work hours								
(9) Part-time	-		-		-0.016*	(0.009)	-0.014***	(0.003)
(10) Full-time	-		-		-0.043***	(0.006)	-0.013***	(0.002)
Sector								
(11) Private	-		-		-0.033***	(0.007)	-0.019***	(0.003)
(12) Public	-		-		-0.010	(0.008)	-0.013***	(0.004)

Note: Figures are estimated coefficients on a non-White dummy variable and on the interaction term UR·Non-White from linear regression models with 1482 time-region fixed-effects. Rows (7) and (8) present results for 'low skill' and 'high skill' non-manual workers, respectively. Also included but not shown are covariates representing age, educational attainment, marital status and children. Standard errors clustered at the time-region level are shown in parentheses. *, ** and *** denote statistical significance at 0.10, 0.05 and 0.01 levels, respectively.

Appendix Table A5: Disaggregated Racial Wage Gaps for Males

	Employed			Log Wage		
	All	Aged 35-64	High Education	All	Aged 35-64	High Education
Black	-0.039***	0.010**	-0.034***	-0.025***	-0.039***	-0.122***
	(0.004)	(0.005)	(0.006)	(0.009)	(0.011)	(0.014)
Asian	-0.124***	-0.135***	-0.073***	-0.031***	0.004	-0.024**
	(0.003)	(0.007)	(0.005)	(0.007)	(0.015)	(0.010)
Other ethnicity	-0.094***	-0.071***	-0.057***	-0.007	0.009	-0.032*
	(0.006)	(0.010)	(0.008)	(0.013)	(0.023)	(0.019)
UR · Black	-0.005***	-0.005***	-0.009***	-0.015***	-0.022***	-0.016***
	(0.001)	(0.002)	(0.002)	(0.003)	(0.005)	(0.005)
UR · Asian	0.004***	0.004	-0.003*	-0.012***	-0.004	-0.023***
	(0.001)	(0.003)	(0.002)	(0.003)	(0.006)	(0.005)
UR · Other	-0.002	-0.010***	-0.001	-0.014***	-0.014	-0.005
	(0.002)	(0.004)	(0.002)	(0.005)	(0.010)	(0.007)
Sample size	2633253	1720654	587095	463550	303669	138961

Note: Figures are estimated coefficients from linear regression models with 1482 time-region fixed-effects. Also included but not shown are covariates representing age, educational attainment, marital status and children. Standard errors clustered at the time-region level are shown in parentheses. *, ** and *** denote statistical significance at 0.10, 0.05 and 0.01 levels, respectively.

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