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Imports, unionization and racial age discrimination in the US

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Past studies of the relationship between competition and racial wages find that domestic competition reduces racial wage discrimination of nonunion workers. This article examines the effects of foreign competition on racial wages of union and nonunion workers utilizing an empirical model which allows for cluster-adjusted SEs by industry. Such a procedure allows independence of observations across industries but not within industries, thereby not overstating the significance of industry invariant controls. In this analysis, clustered SEs prevent the overstatement of the significance of imports as a means to reduce earnings discrimination. We find evidence of a wage premium for nonunion white workers in concentrated industries; however, imports cause the wages of nonunion whites to converge towards market rates. In contrast, for union workers in concentrated industries, wage standardization provides a sanctuary from market power initiated discrimination such that imports play a limited role in reducing discrimination.

I. Introduction

The neoclassical theory of discrimination has its origin in the literature on expense preference (Alchian and Kessel, 1962; Williamson, 1963). The basic premise is that profit in noncompetitive product markets gives managers the latitude to pursue objectives that are utility maximizing rather than profit maximizing (Becker, 1971). Nonprofit maximizing behaviour can result in excessive expenditure on office amenities (Edwards, 1977; Hannan, 1979; Hannan and Mavinga, 1980), workers' wages and employer-initiated wage discrimination (Becker, 1971). Indeed, Becker's theory (1971) postulates that some portion of labour market discrimination is market power initiated. Moreover, the theory suggests that market power initiated discrimination flourishes in profitable, noncompetitive industries, whereas intense pressures to reduce cost in highly competitive industries reduces employers' latitude to engage in discrimination. Further, it follows that competition in either domestic or global product markets reduces industry profit and, therefore, reduces employers' latitude to engage in market power-initiated discrimination.

Recent study by Peoples (1994) tests the relation- ship between domestic competition and racial wage discrimination for union and nonunion workers and finds that increased domestic competition reduces the wage premium for white nonunion workers in manufacturing industries. However, Peoples finds less racial wage disparity for union workers and no evidence that competition induces increased racial wage equality of these workers. He argues that standardized earnings and work rules in union employment protect black workers from market power initiated earnings discrimination (Freeman, 1980). These findings suggest that collective bargaining curbs market power initiated discrimination that has origin in domestic product markets.

Becker's theory indicates that market power initiated discrimination can also be reduced by foreign competition. Agesa and Hamilton (2004) utilize the Public Use Micro

Samples (PUMS) to examine the effects of domestic and global competition on racial wages and find little evidence that imports reduce discrimination. Given Peoples' (1994) findings of a distinctly different impact of domestic competition on the racial wage gap of union and nonunion workers, a limitation of Agesa and Hamilton's study is the lack of controls for worker unionization in the PUMS. Further, their empirical specifications do not allow for a different impact of foreign competition on the wages of union and nonunion workers.

This analysis utilizes a sample of manufacturing workers from the 1995 to 2002 Current Population Survey (CPS) Outgoing Rotation Groups (ORGs) to examine the relationship between global competition and racial wages for workers in high- and low-concentration industries separately by workers' union status. Thus, we examine union and nonunion differences in imports-induced earnings equality when domestic product markets are noncompetitive (highly concentrated industries), relative to the impact when domestic product markets are engaged in fierce domestic competition (less-concentrated industries).

Our empirical examination of the relationship is novel in two dimensions. First, our sub-sample of manufacturing workers from the CPS ORGs omits workers who have imputed wages. Specifically, the wages of these workers are imputed utilizing the earnings of a donor who shares a set of common characteristics. The elimination of imputed earners is necessary, as recent study provides evidence that regressions that utilize imputed earners and include covariates not used in the donor matching criteria (e.g. industry and unionization) suffer severe bias (Hirsch and Schumacher, 2004).

Second, our estimation procedure of the relation- ship utilizes cluster-adjusted SEs, which allow for independence of observations across industries but not within industries (Moulton, 1990; Wooldridge, 2002). Such a procedure prevents the overstatement of the significance of covariates in explaining the dependent variable (Pepper, 2002). In this analysis, clustered SEs prevent the exaggeration of the significance of imports as a means to reduce racial earnings disparity. Moreover, this analysis offers the first study of the relationship between competition (foreign or domestic) and discrimination that takes advantage of the large sample size of the CPS ORGs but does not suffer match bias – additionally providing extensive industry invariant controls utilizing clustered SEs.

II. Background

Effects of imports on wages

We first examine the literature on the effects of imports on the wages of workers in high- and less- concentrated industries. Borjas and Ramey (1995) provide a theoretical model that hypothesizes a negative impact of imports on workers' wages that will be larger for workers in noncompetitive, concentrated industries relative to workers in competitive, less-concentrated industries. Their empirical results support this theory.

Grossman (1984) and Lawrence and Lawrence (1985) provide the theoretical basis for the literature on the effect of imports on union and nonunion wages. Lawrence and Lawrence (1985) present a model in which increased import competition causes an increase in the wages of union workers, whereas Grossman's (1984) model premises on the notions that unions determine wages by majority rule and that they maximize utility of the median worker. In his model, the effects of imports on union wages are ultimately dependent on the elasticity of substitution and therefore, imports could cause an increase or decrease in union wages.

Macpherson and Stewart (1990) provide empirical tests of the effects of imports on union and nonunion wages. Using the May CPS from 1975 to 1981, they find evidence of a significant imports-induced wage penalty for union workers, but no significant penalty for their nonunion counterparts. By contrast, recent study by Shippen and Lynch (2002) re-examines the influence of imports on union wages and find evidence of imports-induced wage penalties for union workers in the period 1983 to 1986, but no significant evidence for the 1987 to 1994 period.

Competition and earnings discrimination

The premise that noncompetitive market structure is associated with increased market power initiated discrimination is a by-product of the expense preference literature. The basic notion is that market power provides employers the opportunity to pursue objectives that are not profit maximizing (Alchian and Kessel, 1962). Thus, in the absence of rigorous product market competition, employers have the latitude to engage in market power initiated discrimination.

Past studies test the notion that domestic competition reduces market power initiated discrimination by single- or multiple-industry approaches. The single-industry approach has been used to examine the effect of deregulation on racial earnings in the motor carrier industry, telecommunications, airlines and rail. With the exception of the airline industry, single-industry studies find convincing evidence that enhanced competition reduces racial earnings discrimination;[1] however, multiple industry analyses have met with mixed success.

Early multiple-industry studies utilize data that delimit industries into imprecise, two-digit industry classifications (Fujii and Trapani, 1978; Johnson, 1978), which may mask the variance from more precisely defined industries, that otherwise could yield significant results.

As a result, these studies find no relationship between domestic competition and market power initiated discrimination.

The next wave of studies utilizes data with more distinct industry classification (Heywood, 1987; Peoples, 1994). Heywood (1987) employs the Panel Study of Income Dynamics (PSID). His wage specification includes racial status, market structure and their interaction to capture the different impact of market structure on black and white wages. The use of an interactive variable to capture racial differences in the impact of market power is the standard in literature (Peoples, 1994; Agesa and Monaco, 2006). He finds strong evidence to support the relationship. Utilizing the CPS, Peoples (1994) examines Becker's theory, emphasizing differences in the relationship for union and nonunion members. He finds no significant effect of market structure on the racial wage gap for union workers and a modest, but consistently significant effect for nonunion workers. His findings

[1] Heywood (1998) provides a comprehensive review of this literature.

support the contention that standardized union earnings protect black workers from market structure driven discrimination (Freeman, 1980).

More recent tests of the market structure/discrimination relationship that allow a different wage structure for union and nonunion workers (separate wage equations) find some support for Becker's theory for nonunion workers (Agesa and Monaco, 2006), whereas studies that restrict the wage determining process to be the same for union and nonunion workers (Agesa and Hamilton, 2004; Coleman, 2004) find no evidence of the relationship. Specifically, Coleman (2004) utilizes a unique dataset which allows the measurement discrimination using respondents' reports of discrimination, wage inequality adjusted for observables, workplace demographics and the probability that the last worker hired will be of a particular race or gender, to test the relationships between market concentration and earnings and employment discrimination of minorities and women. A limitation of this exceptional dataset, however, is the limited overall sample size and the extremely limited number of observations for each subgroup by race and gender, only permitting a single equation estimation of wages for all groups. Such a procedure restricts the wage determining process to be the same for workers by gender, racial and union status. He finds no evidence that domestic competition reduces earnings discrimination by race or gender.

Recent study by Agesa and Hamilton (2004) tests Becker's theory, examining if increased domestic and foreign competition increase racial earnings equality. Taking advantage of the large number of observations in the PUMS, they estimate individual wage equations by industry and race for workers in each three-digit Census Industry Code (CIC) manufacturing industry. Their specification allows the measurement of the impact of domestic and foreign competition on wages discrimination while allowing for industry differences in wage structure. However, the PUMS provides no control for union member- ship. Thus, they do not test for a different impact of imports on racial wages of union and nonunion workers. They find inconclusive evidence that imports reduce discrimination.

The literature on the effects of imports on wages by market structure suggests that workers in concentrated industries experience a larger wage penalty than workers in less-concentrated industries (Borjas and Ramey, 1995). Yet, the outcome of our analysis is not a forgone conclusion, as recent study offers conflicting findings regarding the impact of imports on wages by union status (Macpherson and Stewart, 1990; Shippen and Lynch, 2002). Indeed, in concentrated industries, if imports invoke a large wage penalty on unionized employment – where wage uniformity has been the norm and, consequently, racial earnings disparity is historically less (relative to nonunion employment) – then imports-induced wage penalties for covered workers may be uniformly distributed to black and white covered workers. In contrast, in concentrated industries, if white, uncovered workers benefit more from labour rent sharing relative to their black counterparts, we expect that increased efficiency induced by foreign competition will reduce the wages of uncovered white workers.

III. Data

To test the relationship between global competition and racial wage disparity, we use data from the CPS ORGs for the period 1995 to 2002. We follow the convention of past studies of the relationship, combining multiple years of the CPS to provide a sample large enough to test our hypothesis. Our sub-sample of the CPS ORG contains black and white males aged 16–65 years who are employed full time in manufacturing industries. As individuals appear twice in the ORG files, we eliminate individuals' second appearance in the dataset. We also remove managers from our sample, as they are not routinely covered by collective bargaining agreements. Additionally, roughly 30% of wage earners in the CPS ORG do not report earnings (Hirsch and Schumacher, 2004).[2] Earnings for these workers are imputed using a hot-deck procedure, assigning these individuals the earnings of a donor worker with an identical match for a set of characteristics. A problem occurs because workers' industry and union coverage are not included in the set of matched characteristics. Hirsch and Schumacher (2004) suggest that a match bias problem occurs as a result of including observations with imputed wages in the analyses. Moreover, in this analysis, including workers with imputed earnings results in biased estimates, as the earnings of imputed earners are largely uncorrelated with union coverage, industry concentration and import penetration, key variables in the determination of the impact of foreign and domestic competition on racial earnings. Thus, we eliminate workers with imputed earnings.[3]

To control for industry characteristics, we supplement individual-level data with industry-level data from the 1997 Census of Manufacturers (Bureau of the Census, US Department of Commerce, 1998). Industry characteristics include the capital-to-labour ratio (K/L = industry gross book value of plant and equipment/industry employment), plant size (plant size = industry employment/number of establishments) and four-firm concentration ratios. Industry characteristics from the Census were converted from five-digit North American Industry Classification System (NAICS) to three-digit CIC by weighing industry value of shipments. Additionally, data on union density by three-digit CIC are taken from Hirsch and Macpherson (2003). For each three-digit industry, a 3-year moving average of union density is calculated for each year from 1995 to 2002 to reduce the effect of measurement error in any year.

Data on international trade are from the National Bureau of Economic Research (NBER) US Trade Database, 1989 to 2001 (Feenstra *et al.*, 2002). The impact of trade on an industry is measured by the Import Penetration Ratio (ipr). Particularly ipr for each three-digit CIC industry, in each year 1995 to 2001, is calculated as: ipr = imports/(value of shipments – exports + imports), where the variable, value of shipments for each industry and year is taken from the Census of Manufacturers for that year.

[2] In the 1995 CPS ORGs, imputed wages cannot be identified for workers in the files of January to August. Therefore, all manufacturing workers occurring in these months of 1995 are eliminated from the analysis.

[3] A possible limitation of eliminating imputed earners is that it may create selection bias. Particularly, selection bias of earnings nonresponse by race would undermine the legitimacy of our results regarding the effect of imports on the racial wage gap. To reveal potential nonresponse bias, we examined the proportion of nonresponders in our data by racial group and union status and find roughly equivalent proportions of earnings nonresponders for black and white workers (29% and 31%, respectively), although nonresponse was slightly higher for union workers relative to nonunion (6% points higher). A recent analysis of the bias of earnings nonresponse in the CPS reveals that earnings nonresponse varies little with respect to most earnings attributes but is noticeably highest among workers in the top percentiles of the predicted earnings distribution (Bollinger and Hirsch, 2007), results consistent with our findings of slightly higher nonresponse for union workers, a group that has higher average earnings relative to their nonunion counterparts.

IV. Methodology

We first utilize a switching regression procedure to bifurcate workers into highly- and less-concentrated industries.[4] The critical concentration level is taken to be 50%. Thus, industries with four-firm concentration ratios equal to or exceeding 50% are classified as ‘highly concentrated’, while those lower than this threshold are considered ‘less concentrated’. Highly concentrated industries include engine and turbines, aircraft, ship and boat building, railroad equipment and household appliances. In general, these industries are characterized by high value addition in the production process. In contrast, food industries and other low value-added industries comprise less-concentrated industries. Examples include meat, dairy and bakery products, as well as cement and screw machine products.

There are limitations of using industry concentration as a measure of domestic competition. First, firms in many industries increasingly face competition from firms in previously noncompeting industries; as a result, competitors are increasingly difficult to define within a given industry. Further, because industry concentration is a national statistic, it does not capture local or regional competitiveness. Notwithstanding these limitations, market concentration is the typical measure of domestic competition used in the analyses of industry market concentration and labour market outcomes.

We examine the relationship between global com- petition and racial wages utilizing the following specification:

$$1\text{n}(\text{wage}_{i,\text{un},\text{ms}}) = a + bX + cY + \text{d}Z + f(\text{ipr}) + g(\text{white}) + h(\text{white x ipr}) + u_i \quad (1)$$

$$1\text{n}(\text{wage}_{i,\text{nu},\text{ms}}) = a + bX + cY + \text{d}Z + f(\text{ipr}) + g(\text{white}) + h(\text{white x ipr}) + u_i \quad (2)$$

The variables $1\text{n}(\text{wage}_{i,\text{un},\text{ms}})$ and $1\text{n}(\text{wage}_{i,\text{nu},\text{ms}})$ are the log of real hourly wages of workers (in $2002) of black and white male workers.[5,6] This set of equations is estimated separately by two joint criteria: whether workers are covered or not by collective bargaining agreements (un and nu indicate union and nonunion workers, respectively) and by domestic market structure (ms indicates high- or less-concentrated industries), where X is a matrix of worker characteristics include ages and its square, dummies for marital status

[4] A model was estimated which regressed the natural log of workers’ real hourly wage on their age and its square, marital status, region of residence, union coverage, educational attainment, occupation, unemployment, plant size, the capital–labour ratio and a dummy variable for the concentration ratio. The critical level of domestic competition was obtained by choosing the market concentration level demarcation that maximizes the absolute value of the log likelihood function (Goldfeld and Quandt, 1973).

[5] In the CPS, weekly earnings are top coded at $1923 from 1995 to 1997 and at $2885 from 1998 to 2002. Hirsch and Macpherson (2003) provide estimates of weekly earnings for top-coded males for each year, assuming a Pareto distribution (http://www.unionstats.com/). Hourly earnings of top-coded workers in our dataset are calculated by dividing the Hirsch and Macpherson (2003) estimates of workers weekly earnings by their usual hours worked per week. Additionally, hourly earnings are deflated using the CPI, taken from Economic Report of the President 2003 (Council of Economic Advisors, 2003).

[6] In the CPS, hourly earnings include any overtime pay, commissions or tips usually received. To the extent that overtime is restricted for black workers, our estimates of racial wage disparity in each union status/market structure group provides biased estimates racial wage discrimination for that group.

(married and separated, divorced or widowed with single omitted), six regional dummies (Mid-Atlantic, East North Central, West North Central, South Atlantic, East South Central, West South Central, Mountain Pacific with New England omitted), seven education dummies (middle school, some high school, high school diploma, some college, 2-year degree, bachelor, graduate or professional degree, with elementary omitted) and union coverage. Dummy variables are also included for occupation (technical, administrative, craft, operative transportation and labourers omitted). Industry-level variables include plant size and the capital-labour ratio as well as a 3-year moving average of union density. The controls for time include annual dummy variables. The omitted years are 1995 (which only includes the last 4 months of that year) and 1996. The monthly unemployment rate is also included to control for macroeconomic conditions.

Of particular interest to this study is the impact of import penetration on the wages of blacks and whites in each union status and market structure group. The coefficient on ipr, *f*, captures the marginal impact of a percentage point increase in import penetration on wages of the base group, black workers. If competition mandates that employers are more efficient in the payment of wages, we would expect international competition to negatively impact workers' wages. The coefficient on the interaction term, *h*, captures the differential impact of imports on the wages of whites relative to blacks or, more specifically, the rate at which imports alter the racial wage gap. If increased competition from imports reduces the earnings advantage of whites (relative to their black counterparts in that union status and market structure subgroup), then the coefficient will be negative, indicating that imports increase earnings equality. Because white racial status is nonlinear in our wage estimation (Equations 1 and 2), the effects of white racial status on wages is equal to the partial derivative of hourly wages with respect to white racial status, $\delta \ln(\text{wage}_{i},)/\delta$ (white). The partial derivative of hourly wages with respect to white racial status can be represented as follows:

$$(e^{g} - 1) \text{ x } 100 + (e^{h} - 1) \text{ x } 100 \text{ x ipr} \tag{3}$$

The first term on the right-hand side, $((e^{g} - 1) \text{ x } 100)$, captures the portion of the racial wage gap in a market structure/union coverage group that is independent of imports, where g is the coefficient on white. The second term captures the effect of imports on the racial wage gap, where h is the coefficient on the interactive term; thus, $(e^{h} - 1) \text{ x } 100$ measures the rate at which imports change the racial wage gap in a market structure/union coverage group. We evaluate ipr at the average import penetration ratio for each market structure and union status group to determine the average impact of ipr for that group. It is important to note that our earnings specification assumes that there is no racial sorting into unionized industries. Thus, our estimation of labour market discrimination in each market structure, union status group is limited to gauging wage discrimination.

Recent study highlights that regression estimation that matches group-level data (in our case, industries) to individual-level data (individual workers) biases downward SEs, but not necessarily coefficient estimates (Moulton, 1990; Wooldridge, 2002). The bias is a result of group (industry) effects in the error term. Specifically, Ordinary Least Square (OLS) assumes independence of observations; however, in this analysis, industry characteristics take on the same value for all workers in an industry, violating the assumption of independence. Industry-level variables in our analysis include: plant size, capital–labour ratio, union density as well as key variables, industry concentration, import penetration and the race/ import penetration interaction. As a result, we utilize cluster-adjusted SEs such that observations are assumed independent across industries but not independent within industries. Moreover, OLS SEs overstate the significance of industry invariant controls (Pepper, 2002). In this analysis, the use of OLS SEs would overstate the significance of imports as a means to reduce racial earnings disparity.

The above specification utilizes a single equation to estimate the wages of white and black workers in each union status/market structure subgroup, constraining the structure of earnings to be the same by racial status. Moreover, the tendency for blacks to receive different returns to human capital may distort estimates of racial earnings (Johnson, 1978). As an alternative, we estimate separate wage equations for black and white workers in each subgroup. Specifically,

$$1\text{n}(\text{wage}_{i,r,un,ms}) = a + bX + cY + \text{dZ} + u_{i,r,un,ms} \quad (4)$$

$$1\text{n}(\text{wage}_{i,r,nu,ms}) = a + bX + cY + \text{dZ} + u_{i,r,nu,ms} \quad (5)$$

where the subscript *r* indicates race (black (*b*) or white (*w*)), ms the market structure (high- or less- concentrated industries) and all other variables are as defined in Equations 1 and 2. Clustered SEs are also used in the estimation of Equations 4 and 5.

V. Results

Table 1 presents OLS estimates of racial earnings in each domestic market structure (concentrated and less-concentrated industries), separate for union and nonunion workers. Coefficient estimates for time and regional controls are not presented for brevity. The coefficient on ipr allows the measurement of the effect of global competition on the wages of black workers by union coverage and domestic market structure subgroup. The coefficient on ipr is small and insignificant for each market structure, union status subgroup, indicating that imports impose an insignificant wage penalty for blacks in manufacturing industries (recall, black workers in each subgroup are the base group).

The coefficient on the interaction term reveals the differential impact of imports on white wages relative to the base group of black workers in each union status and market structure subgroup. The interaction term is insignificant for nonunion workers in less-concentrated industries (column 2, Table 1) and for union workers in high- and less-concentrated industries (columns 1 and 3, Table 1, respectively). However, an increase by 1% in import penetration significantly reduces the wages of nonunion whites in high-concentrated industries by an additional 0.68% more relative to black nonunion workers in that domestic market structure group. These findings suggest that imports disproportionately reduce the wages of white nonunion workers in high concentrated industries relative to their black counterparts.[7] These findings support the notion that

[7]The marginal impact of a characteristic on the wage of the group in question is found by taking the exponential of the estimated coefficient minus one and multiplying by 100.

Table 1. OLS estimates of earnings of workers in manufacturing

	Less concentrated		Concentrated	
Variable	Union	Nonunion	Union	Nonunion
Age	0.0339***	0.0553***	0.0333***	0.0507***
	(5.71)	(10.49)	(4.49)	(9.68)
Age^2	-0.0003***	-0.0006***	-0.0003***	-0.0005***
	(-4.84)	(-9.89)	(-3.55)	(-9.08)
Middle school	0.1148***	0.1152***	0.0010	0.2416***
	(3.38)	(3.30)	(0.02)	(5.28)
Some high school	0.1914***	0.1969***	0.0748*	0.1880***
	(4.89)	(9.57)	(1.93)	(3.65)
High school diploma	0.2992***	0.3517***	0.1575***	0.4052***
	(9.34)	(13.34)	(3.70)	(8.96)
Some college	0.3829***	0.4345***	0.2206***	0.5110***
	(10.97)	(14.96)	(5.59)	(8.53)
2-year degree	0.415***	0.4806***	0.2315***	0.5282***
	(11.54)	(15.33)	(5.54)	(11.47)
Bachelor	0.4156***	0.6514***	0.3128***	0.7423***
	(9.88)	(18.07)	(6.27)	(9.52)
Graduate or professional school	0.4823***	0.8744***	0.4695***	0.8921***
	(6.72)	(18.91)	(3.74)	(14.54)
Married	0.1018***	0.1540***	0.0992***	0.1308***
	(5.83)	(10.73)	(9.08)	(6.10)
Separated, divorced or widowed	0.08889***	0.0862***	0.0703***	0.0821***
	(5.18)	(7.30)	(3.31)	(2.44)
Metropolitan statistical area	0.0740***	0.0798***	0.1545***	0.1067***
	(4.21)	(5.89)	(7.60)	(3.73)
Monthly unemployment rate	-0.1944	-0.0210	-0.0152	-0.5823
	(-0.93)	(-1.62)	(-0.79)	(-1.48)
Capital-to-labour ratio	0.0013***	0.0001**	0.0021	0.0001
	(2.30)	(2.63)	(1.35)	(1.52)
Union density	0.4429***	0.2804***	1.0672***	0.3411
	(3.68)	(2.40)	(2.74)	(1.62)
Import penetration ratio	0.0008	0.0004	-0.0051	0.0019
	(0.68)	(0.63)	(-0.83)	(0.51)
White	0.0950***	0.1050***	0.0613	0.2378***
	(3.27)	(6.27)	(0.48)	(5.51)
White x import penetration	0.0004	0.0011	-0.0006	-0.0069***
	(0.63)	(1.64)	(-0.14)	(-3.53)
Constant	9.2309***	8.3186***	9.1010***	8.7705***
	(56.46)	(41.52)	(31.34)	(34.93)
Sample Size	5717	24091	2021	3297
R^2	0.2962	0.4929	0.3480	0.5272

Notes: Each regression also includes controls for year in sample, seven regional dummies, five occupational controls and average plant size.
*, ** and *** denote test statistic significance at the 10, 5 and 1% levels, respectively.

imports induce increased racial earnings equality for non- union workers in concentrated industries.[8]

In each union status and market structure sub- group, the partial derivative of wages with respect to white racial status, $\delta\ \ln(\text{wage}_i)/\delta(\text{white})$, captures the white wage advantage. We evaluate ipr at the average import penetration ratio for each market structure and union status group[9] Table 2 provides a more easily interpretable summary of the two components that make up the partial effects of white racial status for each union status/market structure group. The first term indicates the impact of coefficient g, the portion of the white wage advantage that is independent of foreign competition. If we compare the first term of the partial effects for nonunion workers in high- and less-concentrated industries (26.85% and 11.07% points, respectively), the advantage is much smaller in less-concentrated industries. Our finding

Table 2. The partial effects of white racial status evaluated at average import

	Penetration ratios	
	High-concentrated industries	Less-concentrated industries
Union workers	6.32–1.39=4.93	9.97–0.90=9.07**
Nonunion workers	26.85 - 13.69 =13.13**	11.07þ2.66=13.73***

Notes: Partial effects are evaluated at the average ipr for that market structure, union status group. The first term is calculated: $(e^{g} - 1)$ x 100, where g is defined in Equations 1 and 2. The second term is
calculated as: $(e^{h} - 1)$ x 100 x ipr, where ipr is the average ipr and h is defined in Equations 1 and 2.
F-tests reveal if the above test statistics are significantly different from zero.
** and *** denote test statistic significance at the 95 and 99% levels, respectively.

For nonunion workers that the portion of the white wage advantage that is independent of foreign competition is smaller in less-concentrated industries relative to concentrated industries indicate that fierce domestic competition promotes increased racial earnings equality for nonunion workers (Heywood, 1987; Peoples, 1994). For union workers, the portion of the white wage advantage that is independent of foreign competition is 9.97% and 6.32% points in less- and high-concentrated industries, respectively. Larger racial wage gaps that are not attributable to imports in less-concentrated industries may indicate that collective bargaining is less effective in standardizing wages by race in less-concentrated industries.

In Table 2, the second term of the partial derivative of wages with respect to white racial

[8] It may be that there is selection to union coverage. To test this notion, we employ the Heckman (1976) sample selection procedure. In the first stage, we perform reduced form probit models of union coverage separately for workers in high- and less-concentrated industries. The explanatory variables include controls for race, marital status, region, education, experience, occupation, time, industry and a dummy variable that takes a value of 1 if the worker's state of residence is a right-to-work state and 0 otherwise, for identification (Hirsch and Berger, 1984). From this estimation, the authors generate the inverse mills ratio (lambda) separately for each union status/market concentration subgroup. In a second stage, Equations 1 and 2 were run and that also included the respective inverse of mills ratio for each union status/market concentration subgroup. We find little evidence of selection bias in our estimates of the effect of imports on racial discrimination. Particularly, we find no marked difference in the magnitude or significance of our race, import and interaction coefficients relative to OLS results reported in Table 1

[9] Average ipr is 23.25 and 19.84 for covered and uncovered workers in concentrated industries, respectively, and 22.42 and 24.21 for their respective counterparts in less-concentrated industries.

status captures changes in the white wage advantage as a result of imports. Recall that *h*, the rate at which imports alter the racial wage gap, is significant solely for nonunion workers in high-concentrated industries. Evaluating this term at the average ipr for this group, we find that in high-concentrated industries, imports reduce the wages of white nonunion workers by 13.69% points, resulting in an overall white wage advantage for this group of 13.13% points. These findings provide support for the notion that imports invoke a large and significant wage penalty for white nonunion workers in high-concentrated industries (relative to their black counterparts). Additionally, we find that the second term of the partial derivative is quite small for union workers in high- and less-concentrated industries (-1.39% and -0.9% points, respectively, indicating that imports play a limited role in curtailing market power initiated discrimination of union workers regardless of the level of national competition.

Tables 3 and 4 separately present regression estimates by racial status in each market structure and union status subgroup (Equations 4 and 5). Particularly, we separately examine the estimated impact of imports on the wages of black and white workers in each subgroup. Imports insignificantly impact black and white workers in each category with the exception of white nonunion workers in concentrated industries. Particularly, an increase by 1% in import penetration significantly reduces the wages of nonunion whites in high-concentration industries by 0.46%, whereas imports insignificantly impact their black counterparts and black and white workers in all other market structure/union status subgroup.

It may be that the insignificant effect of imports on the wages of black in a market structure/union status subgroup stems from large SEs, resulting from the limited number of observations of black workers in the subgroup. Thus, it is useful to contrast the magnitude of the coefficients for black and white workers in each subgroup as further indication of whether imports influence the wages of blacks and whites equally. With the exception of nonunion workers in high-concentrated industries, the magnitude of the imports effect is quite uniform for black and white workers in each subgroup. Indeed, in concentrated industries, the marginal impact of imports is roughly an insignificant 0.5% for both black and white union workers (Table 3), indicating that in concentrated industries, imports-induced wage penalties are equally distributed to black and white union workers. Further, we find that imports marginally but insignificantly increases the wages of union and nonunion blacks and whites in less-concentrated industries by 0.1%. These findings provide additional evidence that imports-induced wage penalties are unique to white nonunion workers in concentrated industries. Further, these findings are consistent with the notion that white nonunion workers in concentrated industries disproportionately benefit from labour rent sharing; hence, they bear the brunt of the burden of imports-induced wage penalties.

It is interesting to contrast the findings of our analysis to the findings if imputed earners are included in the data and utilizing an empirical specification with OLS SEs (rather than clustered SEs). In a separate analysis not shown but available on request, Equations 1 and 2 were run for each market structure/union coverage group; however, imputed earners were included in the analysis and the specification utilized OLS SEs. Although the number of observations increased in each market structure/union coverage group as a result on including imputed earners, the overall fit of each model declined dramatically. Particularly, the adjusted R^2 decreased 0.10–0.15 points lower for each

Table 3. OLS estimates of earnings of workers in concentrated industries

	Union		Nonunion	
Variable	White	Black	White	Black
Age	0.0373***	0.0141	0.0515***	0.0316
	(3.98)	(1.10)	(9.37)	(1.60)
Age^2	-0.0004***	-0.0005	0.0005***	-0.0003
	(-3.35)	(-0.32)	(-9.16)	(-1.49)
Middle school	-0.0628	0.1563***	0.2312***	0.4763
	(-0.91)	(7.03)	(4.71)	(1.08)
Some high school	0.012	0.2895***	0.1565***	0.5963
	(0.28)	(6.22)	(2.79)	(1.42)
High school diploma	0.1227**	0.4678***	0.3917***	0.6600
	(2.41)	(8.21)	(8.89)	(1.55)
Some college	0.1848***	0.5792***	0.4949***	0.7756
	(3.52)	(6.71)	(8.04)	(1.83)
2-year degree	0.2013***	0.4731***	0.5077***	0.8588*
	(3.90)	(4.28)	(10.52)	(2.13)
Bachelor	0.2843***	0.5966***	0.7343***	0.9153*
	(4.97)	(4.74)	(9.14)	(2.03)
Graduate or professional school	0.4350***	0.6579*	0.8707***	1.3276***
	(3.67)	(1.92)	(13.77)	(2.94)
Married	0.0871***	0.1477***	0.1335***	0.1320**
	(7.30)	(5.25)	(6.49)	(2.33)
Single, divorced or widowed	0.0590**	0.1709***	0.0758**	0.1396
	(2.54)	(3.20)	(2.20)	(1.70)
Metropolitan statistical area	0.1615***	0.0512	0.0976***	0.1798***
	(7.51)	(0.92)	(3.20)	(3.48)
Monthly unemployment rate	-0.0203	0.0253	-0.0568	-0.0703
	(-1.06)	(0.34)	(-1.43)	(-0.51)
Capital-to-labour ratio	0.0001	0.0001*	-0.0001	0.0001
	(1.08)	(2.00)	(-1.62)	(0.10)
Union density	1.0068**	1.2414	0.2752	0.4491
	(2.68)	(1.22)	(1.26)	(1.11)
Import penetration	-0.0054	-0.0045	-0.0046*	0.0014
	(-1.20)	(-0.55)	(-1.79)	(0.42)
Constant	9.2091***	8.4409***	9.0128***	8.9238***
	(43.69)	(21.84)	(34.30)	(10.51)
Sample size	1773	239	3027	270
R^2	0.3621	0.3592	0.5305	0.4991

Notes: Each regression also includes controls for year in sample, seven regional dummies, five occupational controls and average plant size.
*, ** and *** denote test statistic significance at the 10, 5 and 1% levels, respectively.

market structure/union coverage group relative to the model with omitted imputed earners. Further, the magnitude of the coefficients on many variables is smaller with the inclusion of imputed earners. Particularly, in high-concentrated industries, the impact of imports on the wages of nonunion whites is less than half the magnitude in the model that includes imputed earners, illustrating that in this case, match bias reduces estimates of imports-induced racial wage equality for this group of workers.

In a second stage, Equations 1 and 2 were run for each market structure/union coverage group. Imputed earners were included; however, the empirical specification utilized clustered SEs. The magnitude of the coefficients remained fairly constant, but the SEs increased, particularly for the industry invariant controls. These findings suggest that past studies of the relationship that utilize the CPS suffer from two effects (Agesa and

Monaco, 2006; Peoples, 1994). The inclusion of imputed earners reduces the explanatory power of the models, possibly, also reducing the magnitude and significance of the market power controls (whether it is a measure of domestic or foreign competition). On the other hand, the use of OLS SEs in this study overstates the significance of industry invariant controls, thus overstating the significance of competition in reducing discrimination[10,11]

It is also necessary to reconcile our results with past findings regarding the impact of foreign competition on racial wage equality. The findings of this analysis indicate imports-induced racial earnings equality exclusively for nonunion workers in high-concentrated industries. Agesa and Hamilton (2004) find an insignificant impact of imports on racial earnings disparity for a combined group of union and nonunion of manufacturing workers in concentrated industries. Moreover, in combining union and non- union workers in concentrated industries, wage standardization and less discriminatory wages of union workers may have overshadowed the significant effect of imports in promoting racial wage equality for nonunion workers in Agesa and Hamilton's analysis, thus resulting in their findings of an insignificant imports effect.

IV. Conclusion

This study examines the relationship between global competition and racial wages for union and nonunion workers in manufacturing industries. Our empirical examination of the relationship is novel in two dimensions. First, our sub-sample of manufacturing workers from the CPS ORGs omits workers with imputed wages. The elimination of imputed earners is necessary, as recent study provides evidence of substantial match bias by including imputed earners in analyses in which variables not included in the donor matching criteria are the key determinants of the results (Hirsch and Schumacher, 2004). In our case, union coverage and industry of employment are not included in the matching criteria. As a result, the wages of imputed earners would be largely uncorrelated with union coverage, industry union density, industry concentration and the import

[10]It is possible that the findings of our analysis are contingent upon the level of industry aggregation. Indeed, the progression of the literature on market structure and discrimination reveals increased precision of estimates of the relationship with increased precision in defining industries, moving from two-digit industry coding (Fujii and Trapani, 1978; Johnson, 1978) to three digit (Heywood, 1987). Thus, it is likely that precision of estimates would continue to increase with even finer industry classification. Notwithstanding, the precision of estimates in this analysis is limited to the three-digit industry coding used in the CPS.

[11]A potential shortcoming of this analysis is that cyclical effects in the data may have resulted in the inaccurate measurement of wage discrimination in each market structure/union coverage subgroup. Particularly, wage discrimination may be countercyclical, increasing with downturns in the economy. To test this notion, we calculate the average unemployment rate for the observation period 1995 to 2002 (4.89%) and separate the data for this analysis into two groups, manufacturing workers in above-average unemployment rate years (1995, 1996, 1997 and 2002) and those in below-average unemployment rate years (1998, 1999, 2000 and 2001). Then, the wage specifications outlined in Equations 1 and 2 were run separately for each unemployment rate (above average or below)/union coverage (union or nonunion)/market concentration (high- or less-concentrated) subgroup. We find little evidence of countercyclical effects of wage discrimination. Particularly, the sign and magnitude of the coefficients of the race, imports and interaction terms are quite similar for above- and below-average unemployment rate groups, albeit coefficients are insignificant given the smaller sample size. It is interesting to note that the proportion of workers in the data who are black is smaller in the above-average unemployment rate group relative to the below-average unemployment rate. This is true for both union and nonunion workers, providing some evidence that employment discrimination may be countercyclical for union and nonunion workers.

Table 4. OLS estimates of earnings of workers in less-concentrated industries

	Union		Nonunion	
Variable	White	Black	White	Black
Age	0.0338***	0.0353***	0.0568***	0.0341***
	(5.08)	(3.32)	(10.40)	(6.17)
Age^2	-0.0003***	-0.0003***	-0.0007***	-0.0004***
	(-4.36)	(-2.52)	(-9.76)	(-5.46)
Middle school	0.1169***	0.0696	0.1155***	0.0561
	(3.20)	(0.69)	(3.13)	(1.19)
Some high school	0.2062***	0.0230	0.2029***	0.0363
	(4.89)	(0.20)	(9.47)	(0.81)
High school diploma	0.3104***	0.1605	0.3626***	0.1386***
	(9.00)	(1.74)	(13.09)	(3.40)
Some college	0.3952***	0.2359**	0.4429***	0.2522***
	(10.86)	(2.36)	(15.03)	(4.63)
2-year degree	0.4215***	0.3103***	0.4907***	0.2533***
	(11.30)	(3.08)	(15.18)	(3.75)
Bachelor	0.1396***	0.2039	0.6607***	0.4487***
	(9.04)	(1.65)	(18.18)	(5.62)
Graduate or professional school	0.4835***	0.2646	0.8794***	0.8136***
	(6.64)	(2.09)	(19.00)	(6.85)
Married	0.1050***	0.0726	0.1566***	0.1323***
	(5.19)	(1.51)	(9.95)	(5.98)
Single, divorced or Widowed	0.0943***	0.0371	0.0902***	0.0495
	(4.22)	(0.65)	(6.76)	(1.92)
Metropolitan statistical area	0.0696***	0.1346***	0.0805***	0.0905***
	(4.00)	(3.52)	(5.53)	(4.17)
Monthly unemployment rate	-0.0166	-0.0420	-0.0246	0.0380
	(-0.75)	(-0.63)	(-1.84)	(0.76)
Capital-to-labour ratio	0.0001	0.0001***	0.0001	0.0001***
	(0.19)	(4.87)	(1.65)	(5.81)
Union density	4.2535***	0.6796***	0.2720	0.3805***
	(3.42)	(5.31)	(2.28)	(2.79)
Import penetration	0.0011	0.0015	0.0014	0.0010
	(1.55)	(1.63)	(1.45)	(1.33)
Constant	9.3061***	9.3725***	8.4010***	8.6332***
	(55.27)	(24.92)	(38.58)	(25.67)
Sample size	5172	545	22332	1759
R^2	0.2934	0.3290	0.4930	0.4048

Notes: Each regression also includes controls for year in sample, seven regional dummies, five occupational controls and average plant size.
** and *** denote test statistic significance at the 5 and 1% levels, respectively.

penetration ratio.

Second, we utilize clustered SEs in our estimation procedure, which allows for independence of observations across industries but not within industries. Such a procedure prevents the overstatement of the significance of covariates in explaining the dependent variable (Pepper, 2002). In this analysis, clustered SEs prevent the exaggeration of the significance of imports as a means to reduce racial earnings disparity.

We find little evidence that imports increase racial earnings equality for nonunion workers in less-concentrated industries. These findings are consistent with the notion that fierce domestic competition invokes increased efficiency and racial wage equity such that imports have a limited role in reducing racial wage discrepancies. Yet, in high-concentrated industries, we find that nonunion whites receive a substantial wage premium that is independent of imports relative to their black counterparts. However, imports cause the wages of this group of nonunion whites to converge towards market rates. These findings indicate that imports promote increased racial earnings equality primarily of non-union workers in concentrated industries. We find no evidence of imports-induced earnings equality for union workers in concentrated industries. Indeed, wage penalties for union workers are quite uniformly distributed to black and white covered workers. These findings suggest that collective bargaining provides black covered workers a sanctuary from market power initiated discrimination.

Two caveats are in order. First, this analysis utilizes data on manufacturing workers to examine the global competition/racial wages relationship; however, the manufacturing sector constitutes a small and declining portion of the US workforce, currently about 15%. As a result, this analysis provides at best an incomplete picture of the effects of competition on racial earnings for all workers in the labour market.

Second, it is important to note that our findings do not indicate that nonunion blacks in concentrated industries are the only blacks who experience earnings discrimination. Indeed, we find that blacks in three of the four union coverage/market structure subgroups face substantial and significant wage disparity that is insignificantly reduced by foreign competition, indicating the limitations of domestic and international competition as tools for mitigating racial wage inequities. Thus, if our findings for manufacturing workers are indicative of the entire labour force, it is likely that overall racial wage equality is not attainable by union coverage or increased domestic and global competition. To this extent, increased government intervention may be a necessary policy prescription.

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