FOREIGN COMPETITION, MARKET POWER, AND WAGE INEQUALITY*

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This paper investigates the link between the trend in the returns to education and foreign competition in concentrated industries. We argue that the impact of foreign competition on the relative wages of less skilled workers depends on the market structure of the industry penetrated. The empirical evidence indicates that employment changes in a small group of trade-impacted concentrated industries can explain not only part of the aggregate rise in wage inequality in the United States, but also some of the differences in the trends in wage inequality across metropolitan areas.

A great deal of recent research in labor economics attempts to document and analyze the huge increase in wage inequality that occurred in the 1980s. The basic facts, documented in Bluestone and Harrison [1988], Katz and Murphy [1992], Levy and Murnane [1992], and Murphy and Welch [1992], are clear: wage inequality and the returns to skills rose substantially beginning in the late 1970s. In particular, there was a sizable increase in the wage ratio between highly educated and less educated workers.

Although the by-now voluminous literature agrees on these facts, there is much less consensus on the causes of the increase in wage inequality. In an early phase of their work, for instance, Murphy and Welch [1989] argued that the aging of the baby boom, and the resulting decline in the number of new college entrants entering the labor market, may be responsible for the increase in the wage premium accruing to college graduates. Another strand of the literature [Bluestone and Harrison 1988] argues that the changing industrial mix of the U. S. economy, particularly the shift away from the manufacturing sector and toward service industries, may be partly responsible for the trends. Other researchers argue that the deunionization of the American economy or the decline in the real minimum wage over the 1980s removed the "safety net" supporting the unskilled wage level [Freeman 1993; Blackburn, Bloom, and Freeman 1990]. Still others argue that the increasing

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internationalization of the U. S. economy, either through international trade [Murphy and Welch 1991; Johnson and Stafford 1993] or through immigration [Borjas, Freeman, and Katz 1992] may account for the secular trend in wage inequality. Finally, even after accounting for all these factors, a large fraction of the increase remains unexplained, and hence some researchers invoke "skill-biased technological change" as the key factor underlying the secular trend in wage inequality [Bound and Johnson 1992; Davis and Haltiwanger 1991; Mincer 1993; Lawrence and Slaughter 1993].

In Borjas and Ramey [1994] we showed that the trade deficit in durable goods and the average log wage differential between college graduates and less educated workers share the same trend. To illustrate the similarities in patterns of the two series over the last 40 years, Figure I graphs the trade deficit in durable goods as a percent of GDP against the log wage differential between college graduates and high school dropouts, standardized for experience, and against the log wage differential between college graduates and high school graduates.[1] The data show that the trade deficit in durable goods as a percent of GDP hits peaks and troughs at the same time as the wage series. The only deviation between the two series occurs during the last few years when wage inequality continued to rise, while the trade deficit improved. Murphy and Welch [1992] were the first to note the similarities in patterns, but argued that durable goods trade was only one of several explanations. The graph shows that the trade deficit alone seems to do very well in tracking the movements in the returns to skills, which is the measure of wage inequality that we will use in this paper. The same

1. The standardized wage differential among the various education groups is calculated from the 1964–1991 Annual Demographic Files of the Current Population Surveys, as well as the 1950 and 1960 Public Use Samples of the U. S. Census. In each of these cross-section surveys, we restrict the calculation to the sample of working men aged 18–64, who worked full-time in the civilian sector in the year prior to the survey, and who were not self-employed or working without pay. A worker is employed full-time if he works at least 48 weeks a year and at least 30 hours per week. We calculated the standardized wage differential in weekly earnings by estimating the following regression model in each of the cross sections:

$$\log w_i = X_i\alpha + \beta \text{ (vector of dummies indexing educational attainment)} + \epsilon_i,$$

where X_i includes a fourth-order polynomial in the worker's age, and the vector of dummies indexing educational attainment includes three variables indicating whether the worker is a high school dropout; whether the worker is a high school graduate; or whether the worker has some college. The omitted variable indicates whether the worker is a college graduate. The log wage differential between college graduates and the respective education group is given by the negative of the coefficient in the vector β. Imports, exports, and GDP are in 1987 dollars and are drawn from CITIBASE.

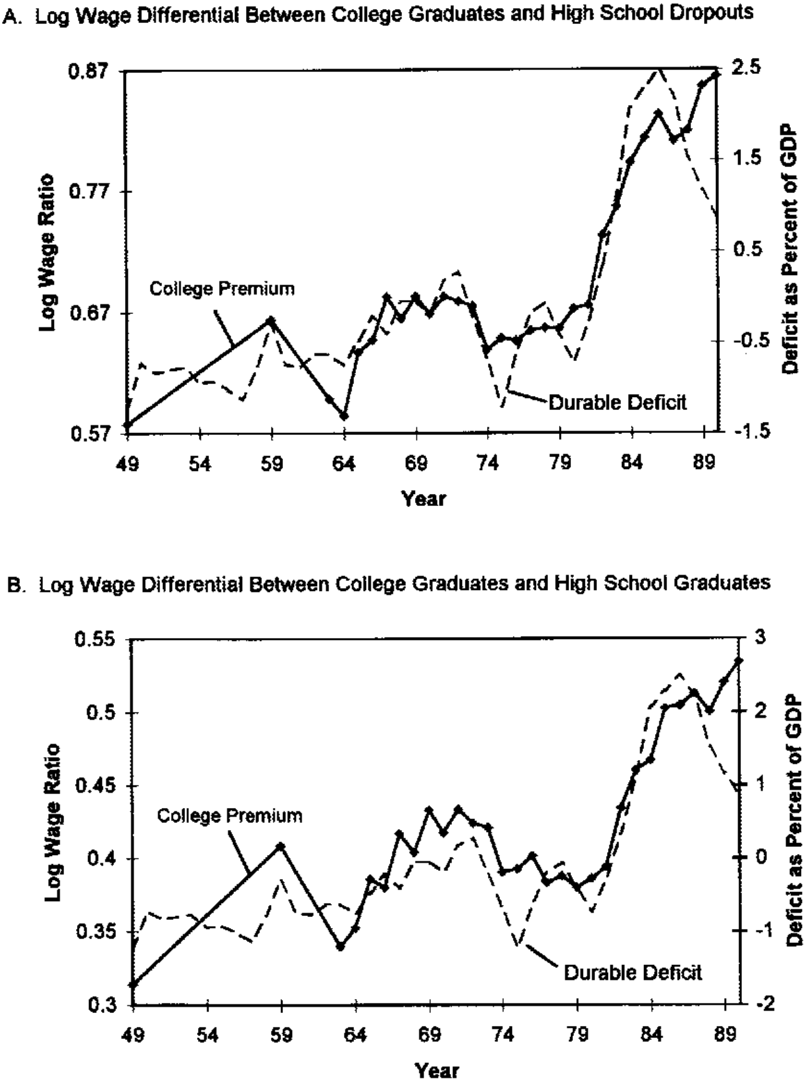

FIGURE I
Durable Goods Trade Deficit and the Returns to Skills, 1949–1990

statement is not true for the trade deficit in nondurable goods. As shown in our earlier paper, the nondurable good trade deficit also rose during the 1980s, but its earlier behavior was very different from that of the returns-to-skills series.

In this paper we present a theoretical interpretation of the time series results, and test the theory using a panel data set on relative wages across metropolitan areas. Our main idea may be stated simply: trade in durable goods can have a strong impact on the returns to education because of the structure of the industries that produce durable goods. Many of these industries employ a disproportionate share of less educated workers, are highly concentrated, earn significant rents, and share those rents with their workers by paying them higher-than-average wages. An increase in imports (or equivalently a decrease in exports) lowers the rents of these industries, leading to a decline in the relative wage of less educated workers. To illustrate this idea, we develop a simple two-sector model in which one of the sectors is an oligopoly. We show that the more concentrated is the industry, the greater is the impact of trade on general wage inequality. We use the theory to suggest why import competition in an industry such as automobiles is more deleterious to the wages of the less educated than import competition in an industry such as apparel.

We test the empirical relevance of our theory in two ways. First, we check the possible quantitative importance of our theory by analyzing changes in employment in trade-impacted concentrated industries and their potential impact on wages. Using accounting exercises, we find that the decline in employment in these industries could account for up to 23 percent of the change in wage inequality. Moreover, we can reasonably link about half of the employment decline to trade.

Second, we test our hypothesis using a panel data set on relative wages across local labor markets. As has been noted by Karoly and Klerman [1994], there is substantial variation in the secular trends in wage inequality across regions. Some areas of the country experienced a substantial increase, while other areas experienced little, if any, increase. We find that the fraction of workers employed in trade-impacted concentrated industries in a local labor market has a significant effect on the relative wage of less educated workers in that labor market. Furthermore, we find that foreign competition in the form of immigrant workers has a statistically and economically significant adverse effect on the wages of the less educated. Overall, we conclude that there is evidence for a link between the number of less educated workers employed in concentrated industries and overall wage inequality. The estimated magnitude of the effect, however, is not large enough to explain the striking time series relationship found in our earlier work.

I. A MODEL OF TRADE, MARKET POWER, AND RELATIVE WAGES

A. Overview

A standard competitive equilibrium model of trade predicts that increased imports will raise the returns to education if the traded goods sector uses a higher proportion of less educated workers than the nontraded goods sector. The standard model, however, cannot explain why trade in different goods should have differential impacts on the college premium, when the different goods have roughly the same proportions of less educated workers.

Why does an increase in net imports of durable goods have a greater impact on relative wages? There are two related characteristics of these industries that distinguish them from other industries. First, the industries that produce durable goods tend to be more concentrated and have higher profits than other industries. For example, in 1977 the four-firm concentration ratio was 93 percent for motor vehicles and car bodies and 86 percent for turbines and turbine generators, while it was 8 percent for women's dresses (*1977 Census of Manufactures*). These high concentration ratios tend to manifest themselves in the performance of the industries. Between 1946 and 1973 the average rate of return in the automobile industry was 16 percent, as opposed to 9.2 percent for all manufacturing corporations and 8.1 percent for all corporations [White 1982]. Second, workers in more concentrated industries tend to earn higher wages. Belman and Weiss [1988] report that the elasticity of the wage with respect to concentration is between 0.07 to 0.20. They found that much of the effect operated through unions, but that there was a significant effect that operated independently of unions. Furthermore, the interindustry wage differentials estimated by Krueger and Summers [1987, 1988] suggest that workers employed in durable goods industries have relatively high wages. In 1984 workers employed in motor vehicles and parts earned 24 percent more than the average worker (holding measured skills constant), while workers in apparel earned 12 percent less. Finally, Freeman and Katz [1991] found that the wage responsiveness to demand shocks varies positively with the degree of unionization.

We consider the following story to be a plausible explanation for the link between trade in durable goods and wage inequality: most of the workers in durable goods manufacturing are high school dropouts or high school graduates. These workers tend to share the rents in their industry in the form of wage premiums; workers in industries with larger rents earn a higher premium. When foreign firms enter markets (domestic or foreign) in which domestic firms have substantial market power, they capture rents that would otherwise go to the domestic industry. This entry increases the relative wage of college graduates in two ways. First, because the rents of domestic firms have fallen, the wage premium of workers remaining in those industries decreases. Second, to the extent that foreign competition reduces employment in the concentrated industries, many of the workers must move to the lower paying competitive sectors of the economy. Overall, the wage of less educated workers falls relative to that of college graduates.

B. Model

We formalize this story in the context of a simple model. For a single country we analyze the general equilibrium of the domestic economy, and take net imports as exogenous. To this end, consider an economy with two sectors, 0 and 1, that produces two consumption goods, x_0 and x_1, and that uses two types of labor, educated labor E and less educated labor L. Let us begin by analyzing sector 1, the noncompetitive sector. Suppose that sector 1 has n firms that behave as Cournot oligopolists participating in a symmetric equilibrium. Foreign countries also produce and consume the good produced by this sector. Total demand for the good x_1 by domestic consumers is given by the inverse demand curve,

$$p_1 = \alpha_0 - \alpha_1 x_1, \tag{1}$$

where p_1 is the price of the good relative to the price of the other good in the economy, and α_0 and α_1 are positive parameters.[2]

The inverse demand curve perceived by each domestic firm i is given by

$$p_1 = \alpha_0 - \alpha_1[x_{1i} + (n - 1)x'_1 + m], \tag{2}$$

where x_{1i} is the amount produced by firm i, x'_1 is the amount produced by each other domestic firm, and m is net imports of the

2. Later in this section we specify the consumer optimization problem from which this demand function is derived.

good. The key assumption of the Cournot model is that each firm takes other firms' quantities as given. Net imports, which are assumed to be exogenous, may be positive or negative. We are implicitly assuming that the markup on goods sold abroad is the same as the markup on goods sold domestically, so that a decrease in exports has the same impact on profits as an increase in imports.

To capture the notion that the concentrated sector is a more important employer of less educated workers than of educated workers, we assume that only less educated labor is required to produce x_1. The production function for firm i is specified as

$$x_{1i} = L_{1i}, \tag{3}$$

where L_{1i} is the number of less educated workers employed by firm i in sector 1.

We assume further that each firm bargains with a union over wages and employment. In the Nash bargaining framework the firm and the union jointly maximize rents, where the rents of firm i are given by

$$\text{Rents}_i = p_1 x_{1i} - w_0 L_{1i}, \tag{4}$$

and w_0 is the wage in the competitive sector. The wage w_1 will be set so that the union receives a fraction γ of the rents, and the parameter γ depends on the exact specification of the union's objective function. For simplicity, we assume that γ is constant.[3]

Given maximization of rents by each firm and its workers, we have equilibrium production of x_{1i}:

$$x^*_{1i} = \frac{\alpha_0 - w_0 - \alpha_1 m}{\alpha_1(n+1)}, \tag{5}$$

and equilibrium rents of firm i and its workers:

$$\text{Rents}^*_i = \frac{(\alpha_0 - w_0 - \alpha_1 m)^2}{\alpha_1(n+1)^2}. \tag{6}$$

To obtain the wage received by the workers employed by the firm, we set the rents captured by the workforce, i.e., $(w_1 - w_0)L_{1i}$, equal to the product of γ and the expression for equilibrium rents above.

3. The specification of the interaction between the firm and the union is identical to the one used in Abowd and Lemieux [1991].

Substitution of the equilibrium value of L_{1i} (equal to x^*_{1i}) yields the following expression for w_1:

$$w^*_1 = w_0 + \gamma \frac{\alpha_0 - w_0 - \alpha_1 m}{n + 1}. \tag{7}$$

Under the assumption of a symmetric equilibrium, domestic industry output is simply nx^*_{1i}. Thus, the number of domestic workers employed in industry 1 is given by nL^*_{1i} or

$$L^*_1 = \frac{n}{n + 1} \frac{\alpha_0 - w_0 - \alpha_1 m}{\alpha_1}. \tag{8}$$

Note that the wage w^*_1 is not allocative, in the sense that labor demand does not depend on w_1. This fact implies that the demand for labor in this sector will generally not equal the supply of labor to the sector. We assume that the parameter values are such that supply always exceeds demand.

We now discuss the partial equilibrium effects of net imports on this sector, taking the rest of the economy as given. First, it is clear from equation (7) that an increase in net imports decreases w^*_1. An increase in net imports reduces rents, so the wage premium to workers employed in the concentrated sector declines. Note that the effect is the same whether the increase is due to an increase in imports or to a decline in the oligopoly's exports to world markets. We can also study how this effect differs when the concentration of the industry changes. A natural indicator of market power in our model is n, the number of firms in the oligopoly. If n is unity, the industry is monopolistic. As n increases, the industry becomes more competitive. Market power affects the wage only through its effect on rents per worker. It is easy to see from equation (7) that $\partial^2 w^*_1/(\partial n \partial m) > 0$. In other words, the more competitive the industry (the larger the n), the lower the impact of net imports on rents per worker and hence the smaller is the decline in wages.

It is also useful to discuss the interaction of net imports and concentration on the equilibrium wage bill of sector 1, which is $w^*_1 L^*_1$. The interesting experiment is a change in the degree of concentration without a change in the relative size of the industry. Thus, we must make adjustments so that the industry size does not change when n changes. There are many ways of making this adjustment. The easiest way is to multiply $\partial(w^*_1 L^*_1)/\partial m$ by the factor $(n + 1)/n$ before taking the derivative with respect to n.

When we make the adjustment, we see that

$$\left.\frac{\partial^2(w_1^*L_1^*)}{\partial n \partial m}\right|_{L_1 \text{ invariant to } n} = 2\gamma \frac{\alpha_0 - w_0 - \alpha_1 m}{(n+1)^2} > 0.$$

Thus, the negative impact of net imports on the wage bill is stronger when the industry is more concentrated, holding industry size constant. This partial equilibrium result forms the basis for a similar result in the economywide equilibrium model.

We now briefly sketch the structure of the rest of the economy. Industry 0 is the competitive sector. We assume that the production function for this industry is Cobb-Douglas, with inputs of educated labor E and less educated labor L_0. Maximization of profits gives the following first-order conditions:

$$w_0 = \beta L_0^{\beta-1} E^{1-\beta}, \tag{9}$$

$$w_e = (1-\beta) L_0^{\beta} E^{-\beta}, \tag{10}$$

where w_0 is the wage of less educated workers in this sector, w_e is the wage of educated workers, the price of x_0 has been normalized to one, and β lies between 0 and 1. These conditions simply equate the wage of each type of labor to its marginal product.

To abstract from distributional effects, we assume that there is a representative worker-consumer, who supplies both types of labor.[4] This consumer has the utility function,

$$U = x_0 + \alpha_0 x_1 - \tfrac{1}{2}\alpha_1 x_1^2, \qquad \alpha_0, \alpha_1 > 0;$$

and labor endowments,

$$L_0 + L_1 \leq 1, \tag{11}$$

$$E \leq \bar{E}. \tag{12}$$

Given prices, it is easy to show that the consumer's demand function for x_1 is exactly as given in equation (1). Because the consumer has no disutility of labor, he will supply his labor inelastically, so equations (11) and (12) will hold with equality. The amount of labor supplied to sector 1 will be determined by the demand for labor in that sector, rather than by the consumer's choice because of the assumed presence of excess supply in

4. Adding two types of consumers would only complicate the model and would not change the basic results concerning the effect of foreign competition on relative wages.

equilibrium. Thus, the supply of labor to the competitive sector will be the excess of the labor endowment over the amount of labor demanded by the oligopoly. Finally, we assume balanced trade, so that the foreigners spend their earnings from the concentrated sector on the goods from the competitive sector.[5] This assumption does not affect the impact of net imports on wage inequality.

C. Equilibrium Effects of Foreign Competition

Because the system of equations is block recursive, we can determine equilibrium w_0, w_1, w_e, L_0, L_1, and E using only equations (7)–(12). The ratio of the average wage of less educated labor to educated labor is a function of these variables, and can be written as

$$\frac{\text{Average Wage of Less Educated Labor}}{\text{Average Wage of Educated Labor}} = \frac{w_0L_0 + w_1L_1}{w_e} = \left[1 + \frac{w_1L_1}{w_0L_0}\right]\frac{\beta\overline{E}}{1-\beta}.$$

Thus, the wage ratio varies with the wage bill in industry 1 relative to the wage bill in industry 0. To determine how the wage ratio varies with changes in net imports of x_1, we need only determine how the relative wage bills vary.

We first establish the effect of m on equilibrium wages and employment. Comparative static exercise gives the following results:

$$\frac{dL_0^*}{dm} = \left[\frac{n}{n+1}\right]\frac{\alpha_1}{[\alpha_1 + (n/(n+1))\beta(1-\beta)L_0^{\beta-2}E^{1-\beta}]} > 0;$$

$$\frac{dL_1^*}{dm} = -\frac{dL_0^*}{dm} < 0;$$

$$\frac{dw_0^*}{dm} = \beta(\beta-1)L_0^{\beta-2}E^{1-\beta}\frac{dL_0^*}{dm} < 0;$$

$$\frac{dw_1^*}{dm} = \left[1 - \frac{\gamma}{n+1}\right]\frac{dw_0^*}{dm} - \frac{\gamma\alpha_1}{n+1} < 0.$$

Thus, an increase in net imports of the good produced by the oligopoly shifts labor from the oligopoly sector to the competitive

5. In particular, the output of the competitive sector is split between domestic consumption x_0, and exports x_0^f. Foreigners' earnings from exporting good 1 are p_1m, so $x_0^f = p_1m$.

sector, and decreases the wages of the less educated workers in both sectors. Wages decrease in the oligopoly sector because rents have decreased. Wages decrease in the competitive sector because the supply of less educated workers has increased.

Combining the results above, it is easy to show that

$$\frac{d[(w_1^*L_1^*)/(w_0^*L_0^*)]}{dm} < 0.$$

Thus, an increase in net imports of x_1 shrinks the wage bill in sector 1 relative to the wage bill in sector 0, resulting in an overall decrease in the average wage of less educated workers relative to educated workers.

How does the impact of foreign competition on relative wages vary with the amount of market power in the penetrated industry? The answer to this important question is summarized in the following proposition.

PROPOSITION. Consider an industry characterized by a Cournot oligopoly that engages in Nash bargaining with its workers. If the industry is more concentrated (but does not change in size), an exogenous increase in net imports of the good produced by that industry will have a larger negative impact on the economywide average wage of less educated workers relative to educated workers.

The result is easily obtained by showing that

$$\left.\frac{d^2[(w_1^*L_1^*)/(w_0^*L_0^*)]}{dn\,dm}\right|_{L_0,L_1 \text{ invariant to } n} > 0.$$

In deriving this result, we apply the same adjustment factor for industry size as we did in the partial equilibrium analysis above. The positive sign of the cross derivative implies that the negative impact of foreign competition on our measure of wage inequality (i.e., the educational wage differential) is smaller when the industry being penetrated is more competitive.

The model presented in this section supports the intuition in the story we told above. Net imports into concentrated industries capture rents that would otherwise be shared with workers in the form of wage premiums. This effect should be greater in industries with higher rents.

The model, while simple and specific, also raises some very interesting questions. The model suggests that the types of goods that a country imports and exports are important. Katz and Summers [1989] have also made this point, based on differential industry wage premiums for all workers. Our model links those premiums to market power, and specifies that less educated workers are the primary recipients.

We conclude this section by discussing an alternative interpretation of our model. Because we seek to provide a theoretical foundation for the time series evidence, we have focused on trade as the source of shifts in relative demand. Other types of forces that shift demand for the good produced by the concentrated sector, however, would have similar effects on wage inequality. In our model, a shift in preferences, such as a decline in α_0, would have effects similar to an increase in m. In a more general model, any kind of preference shift or changes in technology that affected the demand for concentrated goods would have consequences for the rents received by the concentrated sector, and hence for wage inequality. We note this generalization because our empirical work presented below suggests that trade is not the only factor that contributed to the decline of some concentrated industries.

II. Trade, Employment, and Wages in Concentrated Industries

In this section we show that employment changes in a relatively small number of concentrated industries may have had a substantial impact on aggregate wage inequality during the 1980s, and that a significant portion of the employment decline in these industries can be linked to foreign trade.

A. Data

To illustrate the link between employment in concentrated industries, net imports, and wage inequality, we must unavoidably combine a variety of data sources. Our data on wages and employment by education level and industry are drawn from the 1977–1991 Annual Demographic Files of the Current Population Surveys (CPS). The 1976–1990 sample period is selected because the 1977 CPS is the first that identifies a relatively large number of metropolitan areas (which we use in our study of local labor markets in the next section), and because the way that a worker's educational attainment is measured changed substantially beginning with the 1992 CPS. We restrict our analysis to the sample of workers (both men and women) aged 18–64 who worked full-time in the civilian sector in the year prior to the survey, and who were not self-employed or working without pay.[6]

Throughout the study the wage variable is the natural logarithm of average weekly earnings in the calendar year prior to the survey. All wages were converted to 1982 dollars using the GNP implicit deflator for personal consumption. We exclude workers whose weekly earnings fall below $67, and we recoded the wage measure for those workers whose earnings were topcoded by multiplying the topcode value times 1.45. These refinements of the wage data match those used by Katz and Murphy [1992] in their comprehensive study of the wage structure.

We use the worker's completed educational attainment to categorize workers into one of four skill groups: (1) workers with less than a high school education (or "high school dropouts"); (2) high school graduates; (3) workers with some college education; and (4) workers with at least a college degree. To obtain age-adjusted wage differentials across these education groups, we estimated the following regression model in each of the CPS cross sections:

$$\log w_{it} = X_{it}\beta_t + \sum_{j=1}^{4} \gamma_{jt} S_{ijt} + \epsilon_{it}, \tag{13}$$

where w_{it} is the log weekly earnings of person i in calendar year t; X is a vector of standardizing variables; and S_{ijt} ($j = 1, \ldots, 4$) is a vector of dummy variables indicating the worker's educational attainment, where the j index corresponds to the education categories defined above. The vector of standardizing variables X includes a fourth-order polynomial in the worker's age, a dummy variable indicating the worker's gender, and a dummy variable indicating whether the worker is white or nonwhite.

We use the vector of regression coefficients γ from equation (13) to calculate two measures of the age-adjusted returns to skills in each calendar year. In particular, we calculate the standardized log wage differential between college graduates and high school graduates (w_{hs}), and the standardized log wage differential between college graduates and high school dropouts (w_{dp}). These wage

6. A worker is classified as working full-time if he or she works more than 48 weeks during the year, and the usual workweek lasts at least 30 hours.

differentials are given by

$$w_{hs,t} = \gamma_{4t} - \gamma_{2t}, \tag{14a}$$

$$w_{dp,t} = \gamma_{4t} - \gamma_{1t}. \tag{14b}$$

There does not exist a single data set reporting industry employment, shipments, price deflators, and trade for the 1976–1990 period. The National Bureau of Economic Research Immigration, Trade, and Labor Markets (ITLM) data file contains a consistent time series of all of these variables through 1985 covering 450 four-digit manufacturing industries. Obtaining 1990 values for the industry shipment and employment variables was easy because ITLM data are drawn directly from the *Annual Survey of Manufactures*. We used CITIBASE to update the price deflator data. Finally, we used the *Trade and Employment* reports, published jointly by the Bureau of the Census and the Bureau of Labor Statistics, to update the export and import data.[7] In several cases, the trade data reported by the ITLM for the overlapping years differed substantially from the Census-BLS trade data.[8] As a result, we estimate the impact of trade on employment separately for the periods 1976–1985 and 1985–1990.

B. Selection of Trade-Impacted Concentrated Industries

We begin by isolating a group of highly concentrated industries that experienced a significant change in their trade positions over the period. The selection of these industries is complicated by the fact that the concentration ratios, employment, shipments, and trade data are reported for four-digit SIC industries, while the wage and educational composition data are drawn from the CPS, which uses the three-digit Census industry classifications (CIC).

We include any CIC manufacturing industry in our group of industries if (1) the majority of workers in the (CIC) industry were in four-digit (SIC) industries that had four-firm concentration ratios greater than 40 percent (as reported by the 1977 *Census of*

7. *Trade and Employment* reports both quarterly and annual import and export by four-digit SIC code (in most cases). The import data are from "Table 1: Imports for Consumption by Two- and Four-Digit SIC-Based Commodity Groups" and the export data are from "Table 5: Domestic Exports by Two- and Four-Digit SIC-Based Commodity Groups." The reports used are Second Quarter 1987 for the 1985 data and First Quarter 1991 for the 1990 data.

8. The most extreme case for the industries we studied was for nonferrous metals. Imports of nonferrous metals in 1985 were $18,279 million according to the ITLM data (based on aggregating the four-digit industries to the three-digit Census code level), and $10,357 million according to the *Trade and Employment* reports. On the other hand, imports of motor vehicles and parts differed only by 1 percent between the two sources.

Manufactures); (2) the employment ratios in 1976 of high school dropouts to college graduates and high school graduates to college graduates were greater than for the economy as a whole; and (3) the absolute value of the change in net imports from 1976 to 1985 was greater than \$350 million, in 1976 constant dollars.[9] Based on this definition, we included the following industries in the trade-impacted concentrated group: glass (1970 CIC code of 119), primary iron and steel (139, 147), primary nonferrous metals (148, 149), engines and turbines (177), farm equipment (178), construction, mining and materials moving equipment (179), household appliances (199), motor vehicles and parts (219), and ships and boats (228).[10] In 1976 these industries represented 5.5 percent of employment in our CPS sample, and high school dropouts and high school graduates constituted 78 percent of the workers in these industries. In contrast, high school dropouts and high school graduates constitute only 62 percent of the workers employed outside our set of industries.

C. Employment and Wage Changes

We now present an accounting exercise which shows that employment changes in our sample of trade-impacted concentrated industries can plausibly account for a significant fraction of the change in *aggregate* wage inequality. Figure II shows the percent of the labor force employed in these industries by education in the 1976–1990 period. In 1976 almost 8 percent of high school dropouts were employed by these industries, as compared with only 2.3 percent of college graduates. The data also indicate, however, that the percent of less educated workers employed by these industries declined substantially during the 1980s.

Consider the following expression for the average log wage of a particular education group:

$$w_j = w_{cj} + w_{pj} \cdot f_j, \tag{15}$$

where w_j is the average log wage for workers in the jth education group, w_{cj} is the log wage that this group receives in the less

9. For this selection procedure we were forced to use the change over the 1976 to 1985 period because the ITLM trade data for all four-digit manufacturing industries were available only until 1985. We obtained trade data through 1990 for only the selected industries.

10. Although it met the three criteria, we excluded electrical equipment not elsewhere classified (CIC 208) because it mixed very different industries. We also excluded several other industries because the combination of industries within the CIC codes changed over the period.

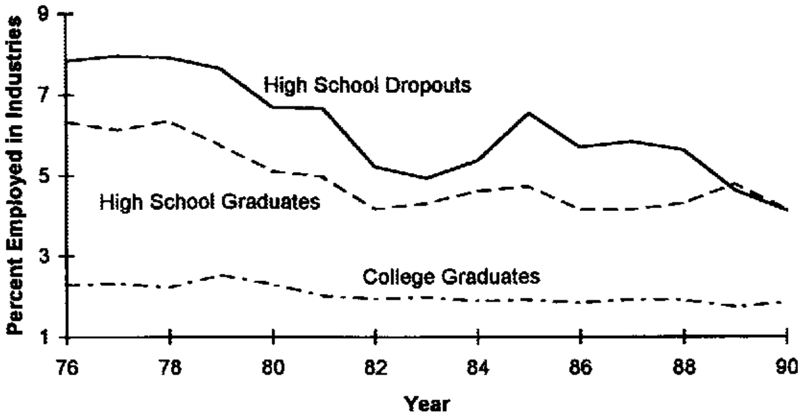

FIGURE II
Employment in Trade-Impacted Concentrated Industries, by Education

concentrated "competitive" sector, w_{pj} is the log of the "industry wage premium" earned by workers in the concentrated sector, and f_j is the fraction of the education group employed in the concentrated sector.[11]

We can decompose changes in the average log wage of the group by totally differentiating equation (15) to obtain

$$\Delta w_j = w_{pj} \cdot \Delta f_j + f_j \cdot \Delta w_{pj} + \Delta w_{cj}. \tag{16}$$

Equation (16) isolates the three factors that can "explain" a change in the average wage of group j: (1) a change in the fraction of workers employed by the concentrated sector; (2) a change in the industry wage premiums; and (3) a change in the competitive sector wage. We use 1976 as the base year for evaluating the equations, meaning that we use the 1976 values of w_{pj} and f_j in evaluating equation (16). We perform the decomposition for each education level, and analyze the differences in the changes of each term for college-educated workers versus high school dropouts and college-educated workers versus high school graduates.

We conduct the decomposition in equation (16) for each of the eleven CIC industries in our group, and then aggregate across industries. The fraction of workers in each education group employed by our industries (as well as by the competitive sector) is

11. For simplicity, we will refer to the less concentrated sector as the competitive sector.

easily obtained from the CPS. We estimated an industry wage premium *for each education level* by regressing the log of the weekly wage (in a particular CPS cross section) on a quartic in age, race, sex, and a vector of eleven industry dummies. In 1976 high school dropouts in our industries earned a weighted average of 26 percent more than high school dropouts outside our industries, high school graduates earned a 17 percent industry premium, and college graduates earned a 16 percent industry premium. Thus, the typical high school dropout benefited relatively more from being employed in the concentrated industries.

Calculating the effect of the change in industrial composition on the competitive wage is somewhat more complicated. We employ a procedure similar to the one used by Borjas, Freeman, and Katz [1992] in their study of the impact of immigration and trade on relative wages. The change in the competitive wage of college graduates relative to, say, high school dropouts that is due to spillovers from the concentrated industry can be written as

$$\Delta(w_{c,cl} - w_{c,dp}) = \eta\Delta[\log(1 - f_{cl}) - \log(1 - f_{dp})], \tag{17}$$

where *cl* denotes college graduate, *dp* denotes high school dropout, and η is the relative factor price elasticity (giving the percentage change in the relative wage of skilled workers for a given percentage change in the relative quantity of skilled workers).[12]

Table I shows the results of the accounting exercise. The actual change in the log of the college wage relative to high school dropouts was 0.191 from 1976 to 1990, and the actual change in the college wage relative to high school graduates was 0.139. Our calculations indicate that the change in the fraction of workers receiving the wage premiums in our group of industries contributed between 0.004 and 0.010 log points, while the change in the industry wage premiums over the 1976–1990 period contributed a negligible amount to the difference between high school dropouts and college graduates and .003 points to the difference between high school graduates and college graduates.

Finally, the spillover of less educated workers to the competi-

12. To see how we derived this expression, note that the overall wage change can be written as $\Delta(w_{c,cl} - w_{c,dp}) = \eta\Delta\log(E_{c,cl}/E_{c,dp})$, where $E_{c,j}$ is the number of type-j workers employed in the competitive sector. The right-hand side can be further decomposed into spillover effects and demographic changes. In particular, $E_{c,j} = (1 - f_j)T_j$, where T_j is the total supply of type-j workers in the economy. Changes in f_j represent industrial composition changes, whereas changes in T_j represent demographic changes.

TABLE I
THE IMPACT OF EMPLOYMENT CHANGES IN TRADE-IMPACTED CONCENTRATED INDUSTRIES ON WAGE INEQUALITY

	College graduates-high school dropouts	College graduates-high school graduates
Actual change in log wage ratio between 1976 and 1990	0.191	0.139
Decomposition		
Due to changing employment in concentrated industries	0.010	0.004
Due to changing wage premium paid by these industries	−0.0006	0.003
Due to spillover effects on competitive sector, assuming relative factor price elasticity = −1	0.035	0.019
Due to spillover effects on competitive sector, assuming relative factor price elasticity = −0.5	0.018	0.010
Due to spillover effects on competitive sector, assuming relative factor price elasticity = −0.32	0.011	0.006

tive sector increased the relative supply of less educated workers so as to raise the college premium substantially. The impact on the wage depends on the relative factor price elasticity. If this elasticity is minus one, then the spillover effect alone generates a 0.035 point change in the log wage ratio of college graduates relative to dropouts and a 0.019 point change in the case of college graduates relative to high school graduates. Borjas, Freeman, and Katz [1992] estimated the relative factor price elasticity between college-educated workers and high school dropouts to be −0.32. This elasticity value would then imply that the employment decline in these industries raised the college premium (relative to high school dropouts) by 0.011 log points.

According to this accounting exercise, the changes in the employment and skill composition of workers employed in our group of industries can explain a reasonable fraction of the decline in the relative wage of less skilled workers over the 1976–1990 period. Most of the impact, however, comes from spillover effects, rather than from composition changes or changes in the industry wage premiums.

D. The Impact of Trade

We chose our group of industries based on their trade experience. We now quantify how much of the change in employment in our group of industries can be linked to changes in net imports. In this exercise we take the change in net imports to be exogenous to supply conditions in the industries, a reasonable assumption if the trade experience of the 1980s was in large part due to the effects of monetary and fiscal policy on the exchange rate (e.g., Sachs [1988]) and to the impact of the oil shocks on the composition of automobile demand. We further assume that the impact of trade on the industry is accurately reflected in changes in the value of shipments.[13] To quantify the full impact of trade on industry demand, we include both *direct* and *indirect* imports and exports. For example, to calculate the impact of trade on employment in the steel industry, we account not only for the direct imports of rolled steel sheets, but also for the quantity of steel sheets imported in the form of fabricated metal, machinery, and transportation equipment. We use input-output tables to account for these indirect effects of trade.[14]

Finally, we convert the effect of dollar-valued trade data into employment. Because total shipments can be decomposed into domestic demand and net imports, we write the following identity for employment:

(18)

$$\text{Employment} \equiv \frac{\text{Employment}}{\text{Shipments}}\,(\text{Domestic Demand} - \text{Net Imports}).$$

If we totally differentiate the expression, we can decompose the changes in employment into changes due to (1) a changing ratio of

13. See Bhagwati [1991] and Deardorff and Hakura [1993] for discussions of the limitations of this approach.

14. We encountered problems in merging the various data sets because the Department of Commerce uses its own set of industry codes for the input-output tables. The only industry that did not match well was ships and boats, so we ignored indirect trade effects on it. There were slight discrepancies between the makeup of the industries in the Department of Commerce codes and the Census codes in the case of primary metals. When input-output tables were used for indirect effects, we aggregated according to the Commerce codes; for direct effects, we aggregated to the Census codes. In any case, the differences were very small. The table used for the input-output analysis is the 1977 Input-Output Commodity-by-Industry Direct Requirements (Table 3) from the May 1984 *Survey of Current Business*. Because many of the matrix elements were near zero and because updating the trade data through 1990 was in some cases difficult, we chose only those columns that had nonnegligible impact on our group of industries. These columns were: 23, 35, 37–55, 57–61.

employment to shipments, (2) a change in domestic demand, and (3) a change in net imports. It is the last change that interests us, so we can write the change in employment due to trade as

(19)

$$\Delta \text{ Employment Due to Trade} = \frac{\text{Employment}}{\text{Shipments}} \cdot (-\Delta \text{ Net Imports}).$$

This expression must be evaluated relative to a base year for the employment-shipments ratio. Because the employment-shipments ratio changed significantly over this period, the choice of the base year may affect the results noticeably. We report full results in the table using 1985 as the base year for the change from 1976 to 1985 and 1990 as the base year for the change from 1985 to 1990. This choice for the base year produces a conservative estimate of the impact of trade on employment because the employment-shipments ratio is lower at the end of the period than at the beginning. We deflate the nominal value of net imports to the relevant base year.[15]

Table II shows the results of the calculations for the two periods, as well as the sum over the entire period. All of the industries in our group experienced substantial declines in employment from 1976 to 1985, as well as for the entire period of 1976 to 1990. Most of the industries experienced a decline in employment from 1985 to 1990, but the aggregate decline during this period was less than 20 percent of the decline during the first period.

Table II indicates that the change in trade explains 68 percent of the total decline in employment between 1976 and 1985. Between 1985 and 1990, however, the trade situation improved for most industries, so that the net effect of trade on employment was positive. If we add the results from the two periods, however, the trade impact is substantially negative. For the entire period the

15. We use industry-specific deflators. The deflators through 1985 are from the NBER database. We used a shipments-weighted average of the four-digit SIC deflators to obtain deflators for the CIC industries. For the period 1985 to 1990 we used the inflation rates implied by CITIBASE price indices to update the NBER indices. We attempted to choose price series that were closest to our industries. The CITIBASE indices used were as follows: glass: the implicit GDP deflator for stone, clay, and glass; primary iron and steel: the producer price index for iron and steel; nonferrous metals: the producer price index for nonferrous metals; engines and turbines: the deflator for investment in engines and turbines; farm equipment: the deflator for investment in farm equipment; household appliances: producer price index for household appliances; motor vehicles and parts: the producer price index for motor vehicles and parts; ships and boats: the producer price index for transportation equipment.

TABLE II
THE IMPACT OF TRADE ON EMPLOYMENT IN HIGHLY CONCENTRATED INDUSTRIES

	1976–1985 employment change (in 1000s)			1985–1990 employment change (in 1000s)			1976–1990 employment Change (in 1000s)	
Industry (1970 CIC Code)	Actual	Due to direct net imports	Due to all net imports	Actual	Due to direct net imports	Due to all net imports	Actual	Due to all net imports
Glass (119)	−33.4	−9.1	−14.8	2.5	2.1	2.4	−30.9	−12.4
Primary iron and steel (139)	−333.2	−29.5	−81.6	0.5	20.2	23.9	−332.7	−57.7
Nonferrous metals (148)	−29.8	−50.5	−81.1	−31.6	26.4	40.1	−61.4	−41.0
Engines and turbines (177)	−25.6	−4.9	−10.3	−15.8	1.0	0.3	−41.4	−10.1
Farm equipment (178)	−57.0	−9.5	−10.2	5.1	−1.7	−1.8	−51.9	−12.0
Construction equipment (179)	−82.5	−62.3	−67.0	−26.3	−4.0	−4.3	−108.8	−71.3
Household appliances (199)	−32.6	−10.4	−10.7	−13.5	5.6	5.8	−46.1	−4.9
Motor vehicles and parts (219)	−46.1	−131.5	−167.7	−46.6	−8.3	−10.5	−92.7	−178.2
Ships and boats (228)	−31.9	−10.6	−10.6	0.2	12.2	12.2	−31.7	1.6
Total	−672.1	−318.4	−454.1	−125.5	53.6	68.1	−797.6	−386.0
Percent of actual change in employment attributable to trade	—	47.4	67.6	—	−42.7	−54.3	—	48.4

decline in employment implied by the changing trade flows equals 48 percent of the total decline in employment.[16]

Overall, trade seems to have been a substantial factor in the employment decline in these industries, but certainly not the only factor. The employment share of trade-impacted concentrated industries fell almost continuously from 1976 to 1990. Much of the fall in employment through 1985 can be linked to trade. The source of the continuing decline in employment after 1985 is not obvious.

Taking the result that trade accounted for half of the employment decline from 1976 to 1990, we can then return to our earlier

16. If we use 1976 as the base year for evaluating the change for both periods, we find that trade accounts for 75 percent of the decline in employment from 1976 to 1985 and for 53 percent of the decline in employment for the entire period. These numbers are slightly higher because they do not allow for technological progress over the period, which reduced the ratio of employment to shipments.

calculation and analyze the overall impact of trade in these industries on the returns to skills. Recall that the impact of employment declines in these industries depend significantly on the assumed relative factor price elasticity. If one believes that this elasticity is as high as minus one, then the combined results suggest that trade could have accounted for a tenth of the rise in the returns to skills. Thus, the trade experience of this small group of industries can explain a noticeable fraction of the rise in the college premium.

III. Evidence from Differences across Metropolitan Areas

A. Motivation

We now test the predictions of our theoretical model using time-series and cross-sectional variation across a number of economies. We wish to test our theory across economies rather than industries because our theory has general equilibrium implications about spillovers into the competitive labor market. Ideally, one would test the theory across countries, but gathering international information on wage inequality and industrial organization is a daunting task. Instead, we test our theory using local economies in the United States. As long as labor is partially immobile across cities, so that relative wage movements differ, we can treat the competitive sector of each city as an isolated economy, at least in the short run. Below, we will show that relative wage movements do vary substantially across cities, suggesting that the assumption of segregated labor markets is reasonable. We would expect that those areas of the country which lose more worker rents would experience a larger increase in the returns to skill, both overall (as a result of composition effects) and in the competitive sector (as a result of spillover effects). We explore these hypotheses by analyzing the determinants of the differences that exist in the wage structure across metropolitan areas.

We have shown that foreign competition was an important factor in the employment decline experienced by our group of concentrated industries. For the purposes of the city analysis, however, competition from other cities within the United States can have an impact as well. If a highly concentrated industry moves from city A to city B, we would expect wage inequality to increase in city A relative to city B. For this reason, and because it is impossible to allocate the U. S. trade deficit to different areas of the country, we link wage inequality to the proportion of workers employed in trade-impacted concentrated industries in the city, as this number is probably the best available indicator of rents flowing to workers within a city.

We use data from the 1977–1991 March CPS files, but we limit the sample to workers living in one of the 44 metropolitan areas that can be identified in each of the cross sections. To obtain the city-specific age-adjusted wage differentials across education groups, we estimated the following regression model in each cross section:

$$\log w_{it} = X_{it}\beta_t + \sum_{k=1}^{44}\sum_{j=1}^{4} \gamma_{jkt}S_{ijt}C_{ikt} + \epsilon_{it}, \tag{20}$$

where w_{it} is the log weekly earnings of person i in calendar year t; X is the same vector of standardizing variables used in the last section; S_{ijt} ($j = 1, \ldots, 4$) is a vector of dummy variables indicating the worker's educational attainment, where the j index corresponds to the education categories defined earlier; and C_{ikt} is a vector of dummy variables indicating the city of residence, with k indexing the metropolitan areas. The regression coefficients γ from equation (20) are used to calculate the two measures of the age-adjusted returns to skills for each locality in each calendar year: the standardized log wage differential between college graduates and high school graduates, and the standardized log wage differential between college graduates and high school dropouts.

It is well-known that there are substantial differences across Census regions in the level and trend of measures of wage inequality (see, for example, Karoly and Klerman [1994] and Topel [1994]). Table III reports the change in the relative wage for the 44 metropolitan areas in our analysis between 1976 and 1990.

It is evident that the phenomenon of increasing returns to skills was not experienced equally by all metropolitan areas. The increase in the returns to skills in some metropolitan areas, such as Los Angeles, is roughly consistent with the aggregate trend. For example, college graduates in Los Angeles earned about 31 percent more than high school graduates and 58 percent more than the high school dropouts in 1976. By 1990 college graduates in Los Angeles earned 47 percent more than high school graduates and 84 percent more than high school dropouts. In contrast, other metropolitan areas, such as Pittsburgh, barely experienced a change in the return to schooling over the period. College graduates in

TABLE III
RETURNS TO SKILLS IN METROPOLITAN AREAS, 1976, 1990

		Log wage ratio of college graduates to:		Percent employed in:		
Region	Year	High school dropouts	High school graduates	Manufacturing	Trade-impacted concentrated industries	Sample size
Entire U. S.	1976	0.549	0.327	28.7	5.5	38,896
	1990	0.740	0.466	22.5	3.3	45,576
Akron, OH	1976	0.403	0.129	37.5	8.2	98
	1990	0.656	0.289	31.2	5.2	130
Albany, NY	1976	0.790	0.517	22.3	6.4	115
	1990	0.517	0.485	13.3	2.4	155
Atlanta, GA	1976	0.800	0.318	20.0	7.2	250
	1990	0.719	0.444	14.6	3.9	283
Baltimore, MD	1976	0.618	0.458	22.5	6.0	406
	1990	0.684	0.582	17.9	3.1	255
Birmingham, AL	1976	0.594	0.332	22.2	9.6	138
	1990	0.851	0.726	14.7	4.2	127
Boston, MA	1976	0.561	0.272	27.5	7.2	376
	1990	0.586	0.414	18.1	5.0	880
Buffalo, NY	1976	0.557	0.343	46.5	16.8	188
	1990	0.385	0.224	31.7	8.0	137
Chicago, IL	1976	0.519	0.350	30.7	5.9	1261
	1990	0.752	0.450	23.2	2.7	1123
Cincinnati, OH	1976	0.415	0.184	37.7	10.3	232
	1990	0.665	0.421	27.7	4.5	276
Cleveland, OH	1976	0.456	0.271	42.9	14.5	364
	1990	0.699	0.415	28.2	6.1	315
Columbus, OH	1976	0.352	0.210	16.6	4.4	147
	1990	0.553	0.380	17.8	2.7	281
Dallas, TX	1976	0.620	0.406	28.6	5.8	322
	1990	0.841	0.474	22.3	2.4	411
Denver, CO	1976	0.458	0.290	19.3	4.0	386
	1990	0.946	0.491	14.9	3.3	297
Detroit, MI	1976	0.538	0.380	41.0	25.5	632
	1990	0.610	0.451	30.7	17.9	918
Fort Worth, TX	1976	0.394	0.299	29.3	11.2	171
	1990	0.745	0.407	17.4	8.6	205
Gary, IN	1976	0.112	0.207	46.4	35.9	94
	1990	0.803	0.583	33.5	15.8	30
Greensboro, NC	1976	0.531	0.296	45.8	0.6	106
	1990	0.569	0.420	36.0	1.2	310
Houston, TX	1976	0.452	0.255	20.5	3.5	455
	1990	0.877	0.513	11.3	1.5	476

TABLE III
(CONTINUED)

Region	Year	Log wage ratio of college graduates to: High school dropouts	High school graduates	Percent employed in: Manufacturing	Trade-impacted concentrated industries	Sample size
Indianapolis, IN	1976	0.532	0.317	29.8	9.7	198
	1990	0.650	0.410	26.7	5.1	69
Kansas City, MO	1976	0.313	0.216	24.6	6.9	244
	1990	0.437	0.340	9.9	1.2	278
Los Angeles, CA	1976	0.578	0.311	29.6	8.5	1320
	1990	0.841	0.471	25.8	5.7	1,793
Miami, FL	1976	0.735	0.535	18.2	1.7	277
	1990	0.846	0.467	12.1	1.1	413
Milwaukee, WI	1976	0.553	0.337	38.0	14.4	254
	1990	0.588	0.563	32.5	5.5	218
Minn.-St. Paul,	1976	0.589	0.446	27.2	4.7	350
MN	1990	0.590	0.299	20.6	3.0	320
Nassau-Suffolk,	1976	0.471	0.281	19.0	5.3	412
NY	1990	0.628	0.405	15.2	4.5	509
New Orleans, LA	1976	0.481	0.292	13.9	4.3	185
	1990	0.815	0.669	5.3	0.0	115
New York, NY	1976	0.660	0.342	20.4	1.5	1357
	1990	0.894	0.505	12.7	0.8	1604
Newark, NJ	1976	0.492	0.355	39.1	5.4	296
	1990	1.100	0.520	24.7	3.0	482
Norfolk, VA	1976	0.745	0.567	19.3	12.2	104
	1990	0.540	0.514	18.8	12.4	147
Patterson-Clifton-	1976	0.688	0.350	35.1	2.9	256
Passaic, NJ	1990	0.845	0.521	26.8	2.7	357
Philadelphia, PA	1976	0.497	0.270	29.6	5.5	695
	1990	0.760	0.482	22.0	3.0	884
Pittsburgh, PA	1976	0.618	0.387	35.1	17.3	385
	1990	0.638	0.468	17.4	6.6	332
Portland, OR	1976	0.428	0.187	18.8	4.5	155
	1990	0.400	0.333	27.7	4.6	192
Rochester, NY	1976	0.275	0.159	45.6	28.2	151
	1990	0.488	0.275	33.8	13.3	206
Sacramento, CA	1976	0.379	0.252	6.7	0.0	137
	1990	0.844	0.531	11.7	2.7	178
San Bernardino,	1976	0.451	0.257	20.2	3.2	189
CA	1990	0.706	0.502	20.5	5.3	261
San Diego, CA	1976	0.571	0.451	20.7	7.2	236
	1990	0.887	0.579	20.5	4.5	228

TABLE III
(CONTINUED)

Region	Year	Log wage ratio of college graduates to: High school dropouts	High school graduates	Percent employed in: Manufacturing	Trade-impacted concentrated industries	Sample size
San Francisco-	1976	0.312	0.210	16.3	4.7	549
Oakland, CA	1990	0.802	0.351	13.6	2.1	463
San Jose, CA	1976	0.558	0.359	36.9	9.1	227
	1990	0.711	0.438	47.4	11.8	189
Santa Ana, CA	1976	0.550	0.376	29.8	9.7	317
	1990	0.771	0.416	26.5	3.9	318
Seattle, WA	1976	0.220	0.196	24.8	12.9	237
	1990	0.532	0.293	24.9	12.5	247
St Louis, MO	1976	0.574	0.351	27.9	8.2	385
	1990	0.790	0.457	20.6	6.3	149
Tampa-St.	1976	0.506	0.301	17.0	1.9	172
Petersburg, FL	1990	0.718	0.523	11.8	2.6	374
Washington, DC	1976	0.711	0.336	5.0	1.4	738
	1990	0.704	0.481	7.2	1.5	904

Pittsburgh earned 39 percent more than high school graduates in 1976, and only 47 percent more in 1990.

B. Empirical Results

The main objective of our empirical analysis is to test whether our theory, which was formulated to explain the aggregate time-series behavior of wages, can also explain the intercity differences in the evolution of the wage structure during the 1980s. In particular, consider the regression model:

$$y_{kt} = I_{kt}\beta_0 + Z_{kt}\beta_1 + \nu_k + \mu_t + \epsilon_{kt}, \tag{21}$$

where the dependent variable y_{kt} is the college premium in metropolitan area k in year t (relative to either high school dropouts or high school graduates); I_{kt} is a vector of variables indicating the industrial composition of the labor market; Z_{kt} is a vector of variables describing other characteristics of the locality, including the fraction of the adult population that is foreign-born, the female labor force participation rate, and the locality's unemployment rate; ν_k is a metropolitan area fixed effect; μ_t is a period fixed effect; and the disturbance ϵ_{kt} is assumed to be uncorrelated with the other variables in the model.[17]

The vector I_{kt} decomposes manufacturing employment into four categories: the proportion of workers employed in our group of trade-impacted concentrated industries; the proportion employed in trade-impacted competitive industries, the proportion employed in other durable goods industries, and the proportion employed in other manufacturing. Our theory suggests that changes in employment in the first of these categories should have the greatest adverse impact on wage inequality. In the last section we defined the sample of industries that were trade-impacted and concentrated. Because a very large fraction of the trade in the *competitive* sector was concentrated in two industries, we used the proportion of workers employed in apparel (151) and footwear (221) to measure the relative importance of the trade-impacted competitive sector. In 1976 high school dropouts and high school graduates constituted 89 percent of their workforce. Thus, these industries have an even higher factor ratio of less educated workers than the high concentration industries.

Columns 1 of Table IV indicate that there is a strong negative correlation between the returns to skills and the proportion of the workforce employed in trade-impacted concentrated industries. For instance, a ten-point increase in the proportion of workers who are employed in these industries increases the relative wage of high school dropouts by 7 percent, and that of high school graduates by 3 percent.

17. The fraction of the adult population in the metropolitan area that is foreign-born is calculated from the 1970 Census, the 1980 Census, and the 1989 CPS. We interpolated the intervening years. The female labor force participation rate (for women aged 18 to 64) is obtained from the "unemployment status recode" variable for the reference week in each CPS. The unemployment rate is also calculated from the employment status recode variable for the reference week and refers to the subsample of men (aged 18 to 64). Note that both of these variables, unlike the earnings variable we use in our analysis, refer to the calendar year of the survey, not to the calendar year prior to the survey. The data used in the regressions are "timed" so that all variables refer to the same calendar year. The 1976 CPS, however, identifies only 32 of the metropolitan areas that we use in our analysis. For the twelve missing metropolitan areas we imputed the values of the female participation rate and unemployment rate. In particular, we assumed that the change in the values between 1976 and 1977 for the metropolitan areas was proportional to the change in the aggregate values. Finally, the fraction of the labor force in the metropolitan area that is unionized, a variable that we use below, is obtained from Curme, Hirsch, and Macpherson [1990] and Hirsch and Macpherson [1993]. The CPS did not collect information on unionization status for the calendar year 1982. We imputed the 1982 value for each metropolitan area by interpolation of the 1981 and 1983 data.

TABLE IV
DETERMINANTS OF LOG WAGE DIFFERENTIAL BETWEEN COLLEGE GRADUATES AND LESS EDUCATED WORKERS

	Wage differential between college graduates and high school dropouts				Wage differential between college graduates and high school graduates			
Variable	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
Proportion in trade-impacted concentrated industries	−0.662 (0.117)	−0.652 (0.112)	−1.045 (0.243)	−0.449 (0.255)	−0.278 (0.082)	−0.293 (0.075)	−0.887 (0.171)	−0.330 (0.174)
Proportion in trade-impacted competitive industries	0.565 (0.759)	1.462 (0.732)	1.344 (0.935)	1.970 (0.916)	0.159 (0.534)	1.019 (0.488)	0.269 (0.661)	0.782 (0.627)
Proportion in other durables	−0.488 (0.079)	−0.366 (0.077)	−0.508 (0.189)	−0.137 (0.194)	−0.223 (0.055)	−0.104 (0.051)	−0.199 (0.133)	0.150 (0.132)
Proportion in other manufacturing	−0.030 (0.153)	−0.015 (0.146)	−0.508 (0.263)	−0.162 (0.261)	0.088 (0.107)	0.092 (0.097)	−0.469 (0.185)	−0.173 (0.177)
Proportion of population foreign-born	0.477 (0.079)	0.331 (0.077)	1.245 (0.185)	0.656 (0.202)	0.141 (0.056)	−0.003 (0.051)	0.489 (0.131)	−0.084 (0.138)
Proportion of women in labor force	0.497 (0.092)	−0.014 (0.109)	0.584 (0.112)	0.021 (0.152)	0.296 (0.065)	−0.198 (0.072)	0.473 (0.079)	−0.071 (0.103)
Unemployment rate	0.104 (0.171)	−0.334 (0.190)	0.382 (0.176)	0.351 (0.205)	0.075 (0.120)	−0.264 (0.126)	0.178 (0.124)	0.227 (0.140)
Includes period fixed effects	No	Yes	No	Yes	No	Yes	No	Yes
Includes metropolitan area fixed effects	No	No	Yes	Yes	No	No	Yes	Yes
R^2	0.253	0.340	0.490	0.533	0.109	0.295	0.384	0.473

Standard errors are in parentheses. The regressions have 660 observations. The observations are weighted by the inverse of the sampling variance of the dependent variable.

The relative importance of employment in trade-impacted concentrated industries as a determinant of the returns to skills is particularly evident in the regressions that include vectors of period fixed effects or city fixed effects. In every case the coefficient on employment in these industries is more negative than for employment in any of the other industries. Employment in the so-called competitive industries always enters with a positive sign. This positive coefficient might be caused by a tautological relationship: these industries pay lower than average wages to unskilled workers. Even when the regressions control for both sets of fixed effects, the returns to skills are still significantly negatively correlated with the proportion of the workforce employed in trade-impacted concentrated industries.[18]

The regressions in Table IV also include a number of control variables that have independent effects on the wage structure. For instance, the proportion of the adult population that is foreign-born has a strong negative impact on the relative wage of high school dropouts. The coefficient is not only statistically but numerically important. Even after controlling for period and city fixed effects, a ten-point increase in the proportion of workers who are foreign-born lowers the wage of high school dropouts (relative to that of college graduates) by 6.6 percent.[19]

Recent work [Grant and Hamermesh 1981; Topel 1994] also suggests that the entry of women into the labor market may have had a negative impact on the wage of unskilled workers. Our regressions indicate that the labor market's female labor force participation rate generally has a positive impact on the college premium as long as period fixed effects are not included in the regression. Removing the aggregate time-variation in the data eliminates much of the correlation between changes in female labor supply and the returns to skills.

Finally, the regressions include the labor market's unemployment rate so as to control for differential impacts of the business cycle. Table IV suggests a positive correlation between the unemployment rate and the returns to skills after controlling for period and city fixed effects. In other words, unskilled workers are worse off during periods of high unemployment.

C. Robustness of the Results

There is an important sense in which these regression results are not surprising. It is well-known that workers in trade-impacted concentrated industries earn more than other workers, so that a labor market with a higher percentage of workers in these industries will necessarily have a higher average wage for its less

18. We also estimated the model including both current and lagged fractions of workers in the concentrated sector to allow for adjustment lags. With few exceptions, the lagged variables were not significant, suggesting that the effects on wages occurred within the year.

19. Some caution is required when using these results to infer the impact of immigrants on the wage of native workers. The CPS does not contain information on the birthplace of workers, so that part of the impact of immigration on the wage structure is probably due to a change in the composition of the workforce. In particular, the entry of relatively unskilled immigrants would reduce the wages of unskilled workers, even if immigrants had no impact on the earnings of natives.

TABLE V
DETERMINANTS OF RETURNS TO SKILLS FOR WORKERS EMPLOYED OUTSIDE TRADE-IMPACTED CONCENTRATED INDUSTRIES

	Wage differential between college graduates and high school dropouts				Wage differential between college graduates and high school graduates			
Variable	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
Proportion in trade-impacted concentrated industries	−0.387 (0.126)	−0.370 (0.121)	−1.272 (0.266)	−0.746 (0.282)	−0.123 (0.084)	−0.132 (0.077)	−0.902 (0.175)	−0.363 (0.179)
Proportion in trade-impacted competitive industries	0.582 (0.806)	1.607 (0.781)	1.481 (1.016)	2.294 (1.007)	0.107 (0.545)	1.040 (0.500)	0.073 (0.676)	0.720 (0.647)
Proportion in other durables	−0.499 (0.084)	−0.377 (0.082)	−0.522 (0.205)	−0.236 (0.212)	−0.230 (0.057)	−0.110 (0.052)	−0.186 (0.136)	0.136 (0.136)
Proportion in other manufacturing	−0.037 (0.163)	−0.039 (0.156)	−0.784 (0.289)	−0.518 (0.289)	0.073 (0.109)	0.071 (0.099)	−0.478 (0.189)	−0.212 (0.183)
Proportion of population foreign-born	0.450 (0.084)	0.293 (0.082)	1.166 (0.200)	0.665 (0.221)	0.122 (0.056)	−0.027 (0.053)	0.514 (0.133)	−0.034 (0.142)
Proportion of women in labor force	0.470 (0.098)	−0.065 (0.117)	0.495 (0.121)	−0.025 (0.167)	0.285 (0.066)	−0.219 (0.074)	0.422 (0.080)	−0.118 (0.106)
Unemployment rate	0.196 (0.182)	−0.258 (0.203)	0.357 (0.191)	0.308 (0.225)	0.055 (0.122)	−0.317 (0.129)	0.153 (0.126)	0.165 (0.144)
Includes period fixed effects	No	Yes	No	Yes	No	Yes	No	Yes
Includes metropolitan area fixed effects	No	No	Yes	Yes	No	No	Yes	Yes
R^2	0.183	0.272	0.417	0.453	0.075	0.262	0.359	0.440

Standard errors are reported in parentheses. The regressions where the dependent variable is the wage ratio between college graduates and high school dropouts have 659 observations. The regressions where the dependent variable is the wage ratio between college graduates and high school graduates have 660 observations. The observations are weighted by the inverse of the sampling variance of the dependent variable.

educated workers. As a result, the negative correlation between the proportion employed in trade-impacted concentrated industries and the college premium can be interpreted as tautological.

It is instructive, therefore, to investigate whether the impact of spillovers is sufficiently important that the negative correlation persists even after we net out this compositional effect. We do this by investigating whether the industrial composition of the workforce has an impact on the relative wage of unskilled workers employed *outside* trade-impacted concentrated industries. Table V presents regressions identical to those reported earlier, except that the returns to skills are estimated on the subsample of workers employed outside the concentrated industries. The results suggest a statistically and economically significant spillover effect. A ten-point increase in the proportion of the workforce that is employed in the concentrated industries increases the relative wage of high school dropouts by 7.5 percent, and that of high school graduates by 3.6 percent, even after controlling for both city and period fixed effects.

These two estimates are more negative than the estimates from the regressions using all workers. This difference is not necessarily an anomaly. Suppose that wages in the unionized sector are set at the national level, with little variation across localities. In this instance, the proportion of workers in the city employed in the concentrated sector should have little effect on the wage of those in the concentrated industry, once national influences are controlled for with period effects. Thus, the regression that includes those workers in the dependent variable can in principle yield a smaller (in absolute value) coefficient.[20]

Our theoretical model suggests that the impact of net imports in concentrated industries on the wage structure works through the fact that these industries are unionized. It is not surprising, therefore, that if the regressions also control for the proportion of the workforce that is a member of a union, the coefficient of the employment share variable is weakened.

Table VI presents the basic set of regressions that include the metropolitan area's unionization rate. As expected, the unionization variable has a negative impact on the college premium in the locality. A ten-point increase in the unionization rate reduces the relative wage of college graduates by one to two percentage points (depending on whether the base is high school graduates or high school dropouts). The inclusion of the unionization variable does not reduce the coefficient on employment in trade-impacted concentrated industries substantially. In particular, a ten-point increase in the share of employment in these industries still increases the relative wage of high school dropouts by 3.7 percentage points, and that of high school graduates by 3.0 percentage points.

20. The direction of the change in the coefficient depends on two opposing effects: including the nationally set wage of the concentrated industry in the dependent variable lowers the magnitude of the regression coefficient, but using the fraction employed in the concentrated sector to form the weighted average raises the magnitude of the coefficient.

TABLE VI

DETERMINANTS OF RETURNS TO SKILLS AFTER CONTROLLING FOR UNIONIZATION RATE IN METROPOLITAN AREA

	Log wage differential between college graduates and high school dropouts				Log wage differential between college graduates and high school graduates			
Variable	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
Proportion in trade-impacted concentrated industries	−0.362	−0.413	−0.684	−0.367	−0.100	−0.169	−0.596	−0.299
	(0.123)	(0.119)	(0.261)	(0.264)	(0.087)	(0.079)	(0.183)	(0.180)
Proportion in trade-impacted competitive industries	0.505	1.332	1.409	1.935	0.122	1.950	0.323	0.769
	(0.737)	(0.717)	(0.926)	(0.916)	(0.523)	(0.482)	(0.653)	(0.627)
Proportion in other durables	−0.423	−0.323	−0.399	−0.129	−0.184	−0.081	−0.110	0.153
	(0.078)	(0.076)	(0.190)	(0.194)	(0.055)	(0.051)	(0.133)	(0.132)
Proportion in other manufacturing	0.100	0.090	−0.367	−0.137	0.166	0.147	−0.355	−0.163
	(0.150)	(0.144)	(0.264)	(0.262)	(0.106)	(0.097)	(0.185)	(0.178)
Proportion of population foreign-born	0.550	0.403	1.146	0.671	0.185	0.035	0.411	−0.077
	(0.078)	(0.077)	(0.185)	(0.202)	(0.055)	(0.052)	(0.131)	(0.139)
Proportion of women in labor force	0.305	−0.131	0.475	0.033	0.182	−0.257	0.387	−0.066
	(0.095)	(0.109)	(0.115)	(0.152)	(0.067)	(0.073)	(0.081)	(0.103)
Unemployment rate	0.331	−0.065	0.395	0.357	0.210	−0.124	0.189	0.230
	(0.170)	(0.193)	(0.174)	(0.205)	(0.120)	(0.129)	(0.122)	(0.140)
Proportion of workforce unionized	−0.528	−0.436	−0.564	−0.205	−0.317	−0.227	−0.450	−0.079
	(0.084)	(0.082)	(0.157)	(0.171)	(0.059)	(0.055)	(0.110)	(0.116)
Includes period fixed effects	No	Yes	No	Yes	No	Yes	No	Yes
Includes metropolitan area fixed effects	No	No	Yes	Yes	No	No	Yes	Yes
R^2	0.296	0.369	0.501	0.483	0.147	0.314	0.401	0.474

Standard errors are reported in parentheses. The regressions have 660 observations. The observations are weighted by the inverse of the sampling variance of the dependent variable.

We conclude this section by analyzing the quantitative importance of our estimates for explaining the rise in aggregate wage inequality. Between 1976 and 1990 the actual change in the log wage ratio between skilled and unskilled workers was 0.191 for high school dropouts and 0.139 for high school graduates. During the same period the fraction of workers employed in trade-impacted concentrated industries fell by 2.2 percentage points. We can use the estimated coefficients on the proportion of workers employed in concentrated industries to calculate the impact on the returns to skills of workers moving from the concentrated industries to nonmanufacturing jobs.[21] If we use the specification that controls for city fixed effects (but not for period fixed effects), the decline in the fraction employed in concentrated industries can account for between 12 and 14 percent of the increase in the returns to skills.[22] Using our earlier result that about half of the decline in employment was a result of trade, we conclude that the regression estimates imply that increases in net imports accounted for between 6 and 7 percent of the aggregate increase in wage inequality.

IV. Summary

This paper presents and tests the hypothesis that foreign competition in highly concentrated industries was an important factor underlying the increase in the returns to skills observed during the 1980s. Our theoretical framework suggests that net imports of goods produced in concentrated markets have a much larger impact on the wage structure than net imports of goods produced in competitive markets. In particular, foreign competition in concentrated industries transfers rents from less educated workers to foreign producers. The more concentrated the domestic industry, the larger is the loss of rents and hence the greater is the decline in the relative wage of less educated workers.

Because our hypothesis singles out a particular sector of the economy as responsible for an important part of the trend in the college premium, we performed simple accounting exercises to determine whether changes in a set of trade-impacted concentrated industries could have a measurable impact on the returns to skills. We found that the shift of workers out of the concentrated sector into the rest of the economy could account for up to 23 percent of

21. Alternatively, we could match the coefficients on the four manufacturing sectors included in the regressions to the actual decline in those sectors. We do not do this because the coefficient on the proportion in trade-impacted competitive industries is large relative to the other coefficients, but at the same time imprecisely estimated (probably because there are small samples of these workers in most cities).

22. Specifications that include city fixed effects are the most appropriate because they net out idiosyncratic differences across cities. It is not clear, however, that controlling for time effects is appropriate in the context of this simulation. In particular, if there are spillovers across labor markets, in the sense that the decline in concentrated employment in city A has an impact on the wage in city B, the period effects are overcontrolling because aggregate employment in these industries also matters.

the increase in wage inequality between 1976 and 1990, but that number was dependent on an assumed relative factor elasticity of one. We also found that trade could explain about half of the decline in employment in these industries from 1976 to 1990. Thus, the changes in trade in these industries could account for up to 10 percent of the aggregate increase in wage inequality.

We then tested our theory by analyzing whether those cities that experienced employment declines in trade-impacted concentrated industries are also the cities that experienced the largest increases in wage inequality. The empirical evidence supported our hypothesis. There is a robust negative correlation between the share of employment in these industries in a labor market and the relative wages of less skilled workers both over time within a city, and in a cross section of cities. However, when we used the regression coefficients to predict the aggregate impact of employment changes on wage inequality, we found that the decline in employment in concentrated industries could explain 12 to 14 percent of the aggregate increase in wage inequality. Again, linking that number back to the proportion of the employment decline attributable to trade, we found that the estimates imply that trade in the concentrated industries can explain about 7 percent of the aggregate increase in the returns to skills.

Overall, we find support for our theory linking trade in concentrated industries and aggregate wage inequality, but not for the notion that this theory can explain most of the increase in wage inequality over the period under study. Our experience is typical of studies in this literature: when one sums up quantifiable effects, they usually fall significantly short of the actual change in wage inequality. The leading alternative explanation for the change in wage inequality is skill-biased technological change. This theory has not yet been subjected to the rigorous quantification undergone by other theories because no one has produced an accurate measure of this technological change. As a result, the question is far from settled. One can argue that skill-biased technological change must be important because observable trade effects are small. However, one can also argue that trade effects may be greater than current estimates because multipliers are much greater than those captured in the analysis. For example, when the concentrated industries were more dominant in the economy, their wage-setting behavior may have had an important impact on the wage-setting behavior of other firms. The decline of these industries during the 1980s may have reduced their influence on the aggregate wages of less educated workers. Thus, foreign competition in industries such as automobiles may have led to increased wage inequality not just by shifting workers from high wage sectors to low wage sectors, but also by changing the wage-setting behavior of the entire economy. This is a question left for future analysis.

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