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LABOR UNIONS AND THE DISTRIBUTION OF WAGES AND EMPLOYMENT IN SOUTH AFRICA

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Abstract

Labor unions are an important economic and political force in South Africa. Inequality in wage rates is among the largest in the world in South Africa, with African and white workers receiving wages that differ by a factor of five. The complex role of unions in closing and creating this wage gap is assessed in this paper. Union membership among Africa male workers is shown to be associated in 1993 with their receiving wages that are 145 percent higher than comparable nonunion workers in the bottom decile of the wage distribution, and 19 percent higher in the top decile of the wage distribution. Quantile regression estimates also indicate the returns to observed productive characteristics of workers, such as education and experience, are larger for nonunion than union workers. If the large union relative wage effect were reduced in half, we estimate employment of African youth, age 16-29, would increase by two percentage points, and their labor force participation rate would also increase substantially.

KEY WORDS: Labor Unions, Wages and Employment, South Africa

Labor unions are an important economic and political force in South Africa. According to surveys by the Bureau of Market Research, 2.5 percent of African workers in urban areas of South Africa were unionized in 1975, 5.5 percent by 1980, and after officially legalizing unions in the early 1980s, 19 percent were by 1985 (Moll, 1993b). The Congress of South African Trade Unions (COSATU) estimates that the number of union members grew from 400,000 in 1985 to 1,205,612 in 1993, amounting to 37 percent of workers in the latter year, although there were still relatively few union workers in agriculture or domestic work (Baskin, 1994). The union share of the labor force is high for a country with the income level of South Africa (Moll, 1996). In the survey that we will analyze, collected in late 1993, one-third of employees are paid-up union members, with the proportion being higher for nonwhites than for whites, and higher for men than for women (Cf. Table 1).[1]

Inequality in wage rates in South Africa is among the largest in the world. African and white workers receive average wages that differ by a factor of five, although part of the gap in wages between the races can be accounted for by differences by race in years of education and location, roughly half according to one assessment (Mwabu and Schultz, 1995). The complex role of unions in closing and creating this wage gap between races in South Africa has not been recently assessed, and earlier only for samples of urban manufacturing workers (Moll, 1993a). Unemployment and nonparticipation in the labor force is substantial among young Africans but remains low among whites, suggesting “underemployment” of Africans could be exacerbated by union-negotiated wage floors that Industrial Councils may enforce. The goal of this paper is to begin a quantitative exploration of the consequences of

South African unions on the distribution of economic welfare. In their search for means to reduce inequality and stimulate growth, the new democratic government of South Africa, which is historically allied with the union movement, must weigh carefully how labor market policies could facilitate the creation of more equal employment opportunities.

Because labor unions traditionally seek to raise the income of workers, they may increase the share of national income received by labor and conversely reduce the share received by capital, and this change in the factor distribution of income may help, at least in the short run, to equalize personal incomes. But because not all workers are union members, unions may also widen wage inequalities between union and nonunion workers. More individuals might also decide to leave the labor force due to distortions in the wage structure associated with unions. On the other hand, the income inequality among union members might narrow, if unions increase wages by a larger proportion among low-wage union workers than among high-wage union workers. The net effect of these countervailing consequences of unions on income distribution are not readily predicted by theory. Within the limitations of a single cross-sectional household survey, we evaluate several of these relationships in South Africa. In particular, we use quantile regression to suggest how wages differ depending on whether or not a worker is a union member. It is beyond the scope of our data to "endogenize" union membership and explain who gets a union job and who does not, or the extent to which unions enhance the productivity of workers with the same observable characteristics, perhaps by giving them "voice" in the workplace. The union wage effects we estimate may thus overstate (or understate) the "rent" or wage distortion associated with the "monopoly" power of unions to restrict the entry of workers.[2]

1. CONCEPTUAL AND MODELING ISSUES

We consider a two sector (i.e. union and uncovered) model to analyze the wage and employment effects of unions, distinguishing between employed union members and all others who could potentially work in the uncovered sector. The standard framework (Johnson and Mieszkowski, 1970; Lewis, 1986; Pencavel, 1991) permits an examination at the individual level of the relationship between union membership on wages and hours, and at the aggregate level of the regional/local labor market the relationship between average union wage effects and the union coverage in that local labor market on a variety of outcomes: employment, unemployment, school enrollments, and hours if working. The local labor market is defined as the Province or "homeland," of which there are 14 in South Africa, as defined under the old regime.[3]

Pencavel (1995) postulates a third sector, in which there are administered wages, straddling the union and competitive wage sectors. He considers workers in the public sector and multinational corporations as belonging to this third administered wage sector in low income countries, because they are effectively required to pay union negotiated (or close to union negotiated) wages to discourage worker turnover and deter unionization of their workforce. Labor market institutions in South Africa may also provide unions with the means to influence nonunion wages.

The South African Industrial Conciliation Act of 1924 provided a legal framework within which white employer associations and white trade unions could voluntarily form Industrial Councils, which negotiated union wages and determined conditions of employment for all racial groups. The agreement reached by these Councils could be extended to all workers and to all firms in an industry, even if they were not parties to the negotiations. It is unclear, however, the extent to which nonunion workers are able to enforce legally these minimum wage floors, particularly by nonwhite workers in small firms (Bendix, 1995). Moreover, the Minister of Labor may occasionally not approve the extension of a Council's rulings to the whole industry. In the 1980's, Africans began to take over the old system of Industrial Councils from the whites, and this centralized industrial wage setting machinery was also supplemented to allow for plant-level negotiations between management and unions. At the time of our survey in late 1993, the two systems for collective bargaining coexisted in South Africa: centralized by industry through Industrial Councils following the Northern European example, and decentralized at the plant level through the new unionized negotiations process. Moll (1996) hypothesizes that the centralized bargaining system is still sufficiently strong to account for the relative scarcity of smaller (lower wage) firms in South Africa, compared with the size distribution of firms in the nonunion sectors in other low- or middle-income countries.

Because the industrial wage floors are more readily extended to and legally enforced on larger firms, and unions may find it easier to organize in larger firms, the size of firm may affect its likelihood of being unionized, and if not unionized, the likelihood of being in the "administered" wage sector distinguished by Pencavel (1995), over which the Industrial Councils are likely to be influential. Unfortunately, our data does not distinguish the size or form of a worker's firm. We therefore must abstract from this intermediate sector for which there is expected to be a positive spillover of union negotiated wages to other employees, either due to an administered wage sector or enforced minimum wage under the Industrial

Councils. Because our data do not distinguish multinational firms or even plant size, both of which might help to characterize employees whose wages might be more likely to follow union-negotiated settlements, our empirical estimates of union wage effects probably understate the effect of unions on wages, on local labor markets, or on industry-wide wage levels and structures. In other words, some nonunion employees are undoubtedly receiving higher wages than they would in the absence of unions, and we cannot determine with any precision who they might be and examine the magnitude of this spillover effect. Controls for industry at the one-digit level may not capture precisely the influence of Industrial Councils but are later included in wage and employment regressions for that purpose. The dummy variables for industry may also embody imperfectly the effect of administered wages in the public sector (see for example public employees are concentrated in professional and personal services as shown in Appendix Table A-1). It may be noted that almost one-half of African men employed by public corporations are already union members, and one-fourth of those employed by the three levels of the government are in a union (Table A-1). The number of African and white workers in union and nonunion jobs appears to be sufficient at all education levels to permit us later to consider how union wage-effects might differ by a workers' education, as confirmed in the cross tabulations of our sample reported in Appendix A-4.

2. MEASUREMENT OF UNION WAGE DIFFERENTIALS

Differences in wages among workers that are not accounted for by observed productive characteristics of workers but are associated with their union membership status are treated as an indicator of wage bargaining power of unions or the distortion in competitive wages that unions cause by restricting entry. As stated at the outset, this measure of union relative wage advantage could be affected by many other factors: (1) compensation for differences in working conditions in or fringe benefits from union and nonunion jobs (Duncan and Stafford, 1980);[4] (2) union organizational efforts that alter the productivity of the unionized labor force and hence the derived demand for union workers (Freeman, 1980); and (3) unobserved compensating variation in worker productivity that employers select through their hiring of workers at above market union wages. In a general equilibrium model, the worker displaced from the union sector by restricted entry may seek employment in the nonunion sector, and thereby put downward pressure on wages in the nonunion sector (Johnson and Mieszkowski, 1970; Mincer, 1976). Conversely, as emphasized by Pencavel (1995) and Lewis (1963) employers in the nonunion sector may raise wages to reduce the likelihood that their workers would unionize. The union-nonunion difference in wages, expressed in logarithms, is our measure of the union relative wage effect.

To clarify how unions affect the entire distribution of wages among union members compared to the distribution among nonunion wage earners, the union conditional effects on the expected log wage are first estimated by ordinary least squares (OLS) (Lewis, 1986), and this effect on the <u>mean</u> wage is then supplemented by quantile regressions, for the <u>median</u> wage earner and other deciles in the distribution of wage residuals among which we report here only the 10th and 90th percentile estimates to save space. By minimizing the sum of absolute deviations of the residuals from the conditional function, we calculate the coefficients for the wage function at each decile of the residuals (Koenker and Bassett, 1978; Chamberlain, 1994; Buchinsky, 1994). Conditional on observed characteristics of workers, the log wage residual variation may be intuitively thought of as unobserved "ability" and "luck." The question explored here is how does union status of workers interact with the residual "ability" in determining wages? Because this residual ability and the covariates such as union status and education may not be independent, the errors in the quantile regressions may be heteroskedastic, and the standard errors of the quantile regression coefficients would then be biased. Therefore, bootstrap estimation of the asymptotic variances of the quantile coefficients are calculated with 20 repetitions (Efron, 1979; Chamberlain, 1994) and are the basis for the reported asymptotic t-ratios reported in this paper. These methods of quantile regression are outlined in an Appendix.

The parameters to standard wage regressions (ordinary least squares) minimize the sum of squared deviations between observed wages and predicted wages based on a linear combination of covariates, and are the best unbiased linear estimates assuming that the errors in the wage equation are independently, identically, and normally distributed. These assumptions that the errors are identically (homoscedastic) and normally distributed are relaxed in the quantile regressions framework. The quantile regressions describe in a less parametric form how the covariate parameters determining wages may in fact vary over all the quantiles of the distribution of residuals, whereas they are assumed to be fixed at all quantiles for OLS estimates and equal to the those estimated for the mean wage. Also the quadratic error minimization criterion underlying OLS is replaced in quantile regression by a criterion that minimizes the sum of the absolute values of the errors. Outlier observations are thus not given as much weight in the quantile estimates, and if the outliers are thought to be generated by such factors as random errors in measurement, we might conclude that quantile regression is more robust to such errors than is OLS because they weight them down. Deaton (1997) emphasizes the failure of OLS to deal with statistical heteroscedasticity which may be confirmed by quantile regressions and then requires correction of standard errors according to procedures White (1980) has developed. Economic hypotheses may also be formulated to suggest interactions between the residuals and the covariates such that the effect of a covariate will differ for individuals depending on their position in the distribution of residuals (e.g. Mwabu and Schultz, 1996). Quantile regression can provide a less parametrically restrictive description of how covariates affect the entire distribution of wages and may thereby shed more light on economic hypotheses regarding the mechanisms generating wage inequality, both within and between groups defined on standard covariates.

First, the standard Mincerian (1974) specification of the wage function is adopted, with education measured in years of schooling completed as a three level spline (primary, secondary, tertiary), postschooling potential years of experience approximated by a quadratic, and a control is included for rural residence. Second, additional controls for industry and region are added to the wage function that may capture the differential power of unions in different industries and regional labor markets with their technologically dictated scale of firm and location-specific natural resource dependence, or the wage compensating variation for differences in union and nonunion job amenities. Third, the wage structure itself is allowed to differ for union and nonunion workers, by introducing interactions with union status, to assess how wage returns to schooling and potential experience vary between union and nonunion workers. Finally, aggregate measures of local union market power are constructed to account for regional spillover effects on utilization of the workforce, and population in the economy.

A more complete model should deal with the process determining which workers are union members, or obtain a job in a unionized firm (Farber, 1983). But we do not have a satisfactory theory and the requisite variables to identify the potentially endogenous nature of union status. Similarly, industry and regional dummies can be interpreted as the outcome of a matching of workers to jobs, which could also suggest that industry and location are endogenous or correlated with the error in the wage function. This concern led to their exclusion from our first specification of the "reduced form" wage function. As with union membership, industry and location are thereafter assumed to be exogenous in this study, although panel or other special data might in the future permit a better analysis of worker heterogeneity with a matching of persons to jobs, by location, industry and union status.

We found relatively few studies of African labor markets and wage structures that helped us to sharpen hypotheses for empirical testing, or could refine the specification of our admittedly descriptive approach to the wage distribution. Knight (1997) considers the political economy of labor relations in Zimbabwe, but says nothing about union wage effects. He does, however, refer to the uneasy relationship between the government and trade unions after independence. Clearly, one motivation for this study is the potential contradiction between the new South African governments' loyalties to its former supporters in the trade union movement that assisted it in gaining power, and the governments' current objective of reducing unemployment and spreading income earning opportunities more equally (Thompson, 1990; Lipton, 1985; Porter, 1976).

3. DATA

Our analysis is of a national probability sample of the South African population collected in late 1993 by the South African Labor and Development Research Unit (SALDRU) at the University of Cape Town in collaboration with the World Bank. A national sample of 43,974 individuals from 9,000 households was drawn randomly from 360 sample clusters and interviewed during the period August through December 1993. The survey instruments and the sampling methods are described in SALDRU (1994). Following the methodology of a previous study by the authors (Mwabu and Schultz, 1995, 1996), statistical corrections for the selective determination of the wage earner sample did not affect noticeably estimates of the wage function parameters, conditional on the working assumption that nonearned income and assets may influence whether a person is a wage earner but do not affect her or his wage offers, i.e. these variables identify the normally distributed maximum likelihood model with sample selection (Heckman, 1979).

The number of employed survey respondents and the proportion who are paid-up union members is reported in Table 1 by the eighth race and gender groups, by one digit industrial classifications. The share of union members is highest in mining, 73 percent, and substantial in manufacturing, 48 percent, and moderate in utilities, transportation, communication and finance, and professional services, 30 -37 percent, and lower but not negligible in wholesale and retail sales, restaurants and hotels, and construction, 26 -24 percent. The fraction of workers who are union members in agriculture and domestic services is much lower, from 5 to 12 percent. The employed population and the fraction union is further disaggregated by type of employment for African and white males in Appendix Table A-1 to identify the union share in public corporations and government, as well as private corporations, and by education in Table A-4.

Table 2 reports the sample statistics for the eight race and gender subpopulations for persons of labor force age, 16 to 65, and for the estimation sample of wage earners subsequently analyzed to assess union relative wage effects. It is noteworthy that Africans have about seven years of education compared with 10-12 for whites, with colored and Indian groups in between. South Africa is relatively urbanized for Africa, but not uniformly across race groups. Two-thirds of the Africans reside in rural areas whereas only one-tenth of the whites and virtually none of the coloreds and Indians live in rural areas. The rate of unemployment among those who report themselves in the labor force (employed or looking for a job in the prior week) is four times larger for Africans and coloreds than for whites, 12-18 versus 3-4 percent for men and women, respectively, and unemployment rates are nearly twice as large for persons age 16-29.[5] In addition, the proportion of the population in the labor force is lower for Africans than whites, 52 versus 85 percent for men, and 32 versus 62 percent for women. A small part of this lower level of participation among Africans could be accounted for by the larger proportion enrolled in school. Recall, however, the much lower levels of school completion by the Africans than whites. Many Africans who are older than usual school-attending age are seeking more education in 1993, possibly because their schools were disrupted and closed during the turmoil of the late 1980s and early 1990s, and possibly because they face poor prospects of finding a job.

As seen in Table 2, black Africans are 78 percent of the labor force aged population, whereas whites are the second largest group representing 12 percent, while colored and

Indians constitute almost 8 and 3 percent, respectively. Our analysis of union relative wage effects concentrates on African and white males to avoid dealing with other factors affecting labor force participation patterns and wages among women, and to reduce the sample size limitations that would arise in any analysis of coloreds and Indians.

4. QUANTILE REGRESSIONS

It has been suggested that unions reduce the inequality of wages among union workers (Freeman, 1970). This may be achieved by reducing skill differentials in wages, such as the proportional "returns" to years of schooling, or reducing the proportional wage gains associated with age or seniority, or decreasing inequality within groups with the same education and experience, or some combination of these. First, we evaluate the net effect of the union status dummy across deciles assuming that the relative wage structure for union and nonunion workers is the same throughout the distribution of residuals. These OLS estimates of the expected effect on the mean of log wages of the conditioning variables are reported in the last column of Tables 3 and 4 for African and white males. The coefficient on union status is .468 for African males, and for white males it is -.051 implying that an African union worker receives a wage that is 60 percent higher (exp (.468) - 1.0) than a nonunion worker with the same observed characteristics. A white union worker receives a wage that is five percent lower than the comparable nonunion worker. The first three columns of these tables report the estimated coefficients for the bottom, median and top deciles of the distribution of wage residuals. They range for union membership for African males from .895 for the lowest decile to .107 for the highest decile. For white males the coefficient on union status also declines monotonically from .188 for the lowest decile to -.270 for the highest decile. White males in the top 70 percent of the distribution of wage residuals (not reported) receive a lower wage if they are in a union job, whereas all deciles of African males in a union job receive a significantly higher wage than observationally comparable nonunion workers. But the log wage gain for African males from the union job is eight times larger for workers in the lowest decile than for African male workers in the highest decile. This strong inverse association between union membership and wage residuals indicates that in both African and white males union membership is associated with reduced wage inequality among union workers as noted for the U.S. by Freeman (1980), Lewis (1986), and others.

Controlling for nine industrial groups in Tables 5 and 6 confirms that a substantial part of the effect of union membership is associated with one digit industrial categories, perhaps reflecting the influence of Industrial Councils or the spillover of administrated wages by industry. The expected OLS log wage differential associated with union status, controlling for industry, is .191 for African males and -.097 for white males. The union log wages advantage for African males controlling for industry ranges from .35 for the lowest decile to .01 for the top decile, and from .14 to -.25 for white males at the bottom and top deciles, respectively.

In Table 7 the education and experience coefficients are allowed to vary for union and nonunion workers to assess whether the reduction in the residual variances in (log) wages among union workers is also associated with modified returns to observable characteristics of African workers. At the median residual African male worker, almost all of the 7.8 percent wage return to a year of primary education received by a nonunion worker is "forfeited" by the union worker, for whom wages are only 1.4 percent higher per year of primary school. At the secondary school level, the 20 percent return to nonunion workers from an additional year of schooling is reduced by 6.8 percent among union workers. At higher education the 31 percent returns to nonunion workers is reduced to 21 percent among union workers. According to these estimates, if a student thought it was likely that he would be employed in a union job, he would have less private incentive to obtain additional schooling, insofar as it would increase his wage.

Union wage structures tend to also exhibit flatter wage profiles with respect to years of postschooling potential job experience. The first year of job experience earns a six percent increase in wages among median nonunion workers, but only 1.5 percent for union workers. However, the wage profile peaks at a later age for union workers than among nonunion workers, at age 50 compared to age 37, respectively, and thus wages increase more gradually for a longer period for union workers and declines thereafter less abruptly.

Table 8 reports the coefficients on the industry dummies and on union status interacted with the industry dummies in the African log wage equation, allowing as in Table 7 for the wage structure returns on schooling and experience to differ for union and nonunion workers (not reported). The effect of union status alone (next to last explanatory variable in Table 8) refers to the union relative wage effect in the omitted industry category. The union relative wage effects in mining are moderate across all deciles, and mining is therefore specified here as the omitted industry category. The effect of being a union member for the median worker is to increase their wages in every other (than mining) industry compared to the nonunion wage in that industry. The nonunion wages are all less than those received by nonunion male African workers in the mining industry. For example, the log wage coefficient for the median worker in manufacturing is -.63 if a nonunion worker and -.10 (i.e. sum relevant coefficients -.63 +.56 -.027) for a union worker, or the union worker earns a wage that is about 53 percent higher than those received by nonunion workers in manufacturing (exp (-.63) -1.0 = 0.53), or about 10 percent less than a nonunion worker in mining (exp (-.10) -1.0 = -.10). Across the deciles of the wage residual distribution, inequality tends to be larger within most industries compared with the excluded category of nonunion (or union) workers in mining. For example, in manufacturing a male African nonunion worker receives a log wage that is .53 lower than in a nonunion worker in mining in the first decile, but only .07 lower in the top decile of the wage residual distribution, for an interdecile range of .46. If he were a union worker, the difference would be -.03 at the lowest decile and +.06 at the highest decile, for an interdecile range of .09. Wage inequality is generally less within the union than within the nonunion sectors of each industry.

The centralized bargaining system in South Africa is expected to reduce within industry union-nonunion wage differentials, but we find that they remain substantial, at least for Africans. For whites, however, the Industrial Councils may mitigate union-nonunion wage differentials, because white plaintiffs have for some time had access to courts to enforce union-negotiated industry-specific wage floors (Bendix, 1995). Employment opportunities for Africans and whites may differ in the nonunion sector, along dimensions such as firm size and capital intensity, which we cannot measure. African nonunion employment may be concentrated in small, low-technology firms, whereas white nonunion employment may occur more often in larger firms using relatively more advanced technologies. Both hypotheses could help explain why African union-nonunion wage differentials are substantial at all quantiles of the wage residual distribution, whereas they are more modest for whites, even when industry effects are allowed as in Table 8.[6]

5. GENERAL EQUILIBRIUM EFFECTS OF UNIONS

Unions are expected to raise the compensation or improve the working conditions of their members compared to what equally productive workers could find in the competitive uncovered sector. As a consequence of the difference between the wages in the union and nonunion sectors, employers in the unionized sector would substitute other factors for the more expensive labor, reducing union employment. Employers of union workers would also have an incentive to reduce the hours of their workers, other things being equal. Workers displaced from union employment or wanting to work more hours would seek employment in the uncovered sector, putting downward pressure on nonunion wages, if the demand and supply of labor competitively clears in this uncovered sector. If the wage declines in the uncovered sector, some of this sector's workers may find home production or leisure activities more attractive, and participation in the labor force could contract. However, the wage setting power of unions cannot persist in the long run unless the firm's product markets possess some monopolistic features.

The effects of union wage differentials depends on the institutional structure of the economy that may differ from country to country and within the same country over time. When the interests of the union and employers coincide, as when both parties stand to gain from government protection from foreign competition, and domestic consumers stand to lose in terms of increased prices of output, the two groups that stand to gain jointly lobby the government for protection to create rents which they then share (e.g., as estimated during the 1970s in Colombia, see Schultz, 1982). Trade distortions of the domestic South African economy are reported to be substantial (Iyengar and Porter, 1990), but we could not find recent estimates of effective protection by industry for South Africa to compare with our estimates of industry specific union wage differentials estimated in Table 8. Empirical methods are therefore needed to assess the distributional consequences of unions and repeated over time.

In the Harris and Todaro (1970) framework, as adapted by Calvo (1978) to the study of unions, a wage differential between the unionized and uncovered sector induces individuals to queue up waiting for the better-paying union jobs. The union relative wage effect or distortion in the labor market thus causes unemployment to increase, until the expected wages (i.e. the wage times the probability of being employed in the sector) are equalized in the two sectors. The distortion in the union sector is responsible for an inefficient increase in unemployment in the distorted sector, presumably because the unemployed gain a private advantage in obtaining the rents from a union job compared with workers who simultaneously hold a job in the uncovered sector. This implication of the Harris-Todaro model is more realistic when the two sectors are geographically separated, which could prevent job search by those employed in the uncovered sector, as in the rural and urban sectors, than when the sectors are adjacent as in the union (or minimum wage covered) and uncovered sectors.

The welfare losses associated with the union relative wage effect are also a function of the magnitude of these labor reallocation responses of employers and workers. More specifically, the greater is the elasticity of employment with respect to the union wage effect, the greater the inefficiency costs associated with a given union relative wage effect. Although any increase in schooling connected to the union relative wage effect may be a harbinger of improved lifecycle employment opportunities for those having obtained more schooling, this adaptation to distortions in the labor market is still an accommodation to a second-best situation. Any increase in schooling thus induced by unions is associated with a welfare loss, unless other policies have already distorted and depressed school enrollment levels below their optimal levels (Mwabu and Schultz, 1995, 1996).

6. INTER-REGIONAL ESTIMATES OF UNION EFFECTS ON TIME ALLOCATION

To estimate empirically these reallocation effects of unions, we divide South Africa into four provinces and ten homeland territories. Under Apartheid nonwhites were not free to move and seek employment outside of their region of registration. Only the Africans are represented in all of these regions, whereas the colored are concentrated in the Cape Province and the Indians in Natal. The whites are distributed across only the four provinces and are less likely than the nonwhites to be strongly affected by the local labor market power of unions. Our analysis of interregional effects of unions is therefore restricted to Africans, although it was repeated for all nonwhite without modifying any of our conclusions.

Regional labor markets in which the log wage effect of union membership is estimated to be larger are expected to be regions where employment would be lower, and these persons who are not employed would then be enumerated in school, unemployment, or nonparticipation. According to the Harris-Todaro-Calvo model the unemployment rate would be larger in regions where union wage effects were larger. If the elasticity of demand for labor in the union sector with respect to the union wage were negative and constant across regional labor markets, we would also expect the labor reallocation effects into nonemployment would be greater in regions where the share of employees in the union sector is larger. Clearly, if only a small fraction of the regional labor market is unionized, only a small fraction of the resident population could be expected to be displaced from union jobs by any given union wage premium. Of course, union wages and shares may be interdependent, but both may affect labor market allocations.

However, the union share of jobs and the union wage effect may be related to other regional labor market characteristics affecting employment opportunities. In particular, we expect that the union share of employment in a region is related to the predominance of large firms and the industrial mix in a region which could itself affect the level of regional wages, due to exogenous demand factors such as the location of natural resources (e.g., mines and ports) and supply factors such as the educational qualifications of the labor force. If the regional level of wages that is derived from these data, adjusting for education and industry but not for union status, is positively correlated with the union share of employment and wage effects, we want to control for this regional level of wages in order reduce omitted variable bias in our estimation of union wage and union share effects on time allocation.

Empirical studies of labor supply behavior note two regularities. The labor force participation elasticities with respect to a person's own wage opportunities tend to be larger for youth than for prime aged adults, and these elasticities tend to be greater for women than for men, particularly married women compared with married men. Therefore, the analysis of employment effects focuses on only the young African population age 16 to 29, stratified by sex, who are expected to exhibit a more elastic labor supply response to their own wage and employment opportunities. Descriptive statistics for this sample, as well as for the older complement, are provided in Appendix Table A-5.

Four columns in Table 9 report estimates of the probability that the individual is employed (2), unemployed (3), enrolled in school and not in the labor force (4), and none of the previous three categories and therefore a "nonparticipant" in the labor force or school (5). These binary mutually exclusive activity outcomes are fit by the logit model to a series of control variables and three characteristics we construct to describe the regional labor market. The first is the difference between the log wage of union and nonunion workers in the region, having controlled a quadratic in age, years of schooling, rural, rural-regional interactions, and industry dummies. The industry effects are included to allow for the fact that some industries are more likely to be unionized than others, perhaps because their production processes imply economies of scale and large firms are, as discussed earlier, more likely to pay union or administered premium wages. The second regional variable is the share of the region's employees in the race, age, gender group that is unionized.[7] The third variable is a regional estimate of the wage level, implied by the regional dummy coefficients in a log wage regression that does not control for the union membership of the worker, but does include, age, age squared, education, rural, rural-regional interactions, and industry dummies. These three variables for the 14 regions (four provinces and ten homelands based on the old regional boundaries) are predicted variables within the regions, and the regressions in Table 9 therefore report asymptotic t-ratios based on Huber (1967) standard errors, corrected for the regional covariance structure in the errors. Appendix Table A-2 reports the predicted values for the explanatory variables by region.

The first response in allocation of labor due to a union being able to increase wages by restricting entry is for unionized employers to reduce the number of hours of work they demand of the more costly union workers. This may take the form of fewer employees or fewer hours or both, depending on how the costs and productivity of employment vary by hours worked, and the importance of fixed costs per worker, and the conditions of the union contract (e.g., Boal and Pencavel, 1994). It is also possible that the labor supply of the workers would also respond to an increase in their union relative wage premium by wanting to work fewer or more hours, depending on whether the (negative) income effect prevailed over the (positive) substitution effect. Conditional on participation in the labor force, the adjustment of hours to wages tends to be often estimated as negative for males, and although it may be positive for females, it tends to be small (e.g., Schultz, 1980). Our expectations are therefore that the combined demand and supply response will be a reduction in hours, conditional on being an employee.

Column 1 in Table 9 reports estimates of the effect of the regional union relative wage effect on the number of hours per day worked by union members. This effect might be expected to be larger when a larger share of the region's labor market is already unionized. Although it would be desirable to correct these estimates for the possible sample selection bias due to analyzing only persons who are employees, it is not clear to us what variable(s) would provide identification for such a selection model, by being an argument in the sample selection decision rule and not also entering into the hours of work equation directly. The union relative wage effect is also included without interacting it with union membership, and is not statistically significant for nonunion workers, and is therefore not reported in Table 9. The hours regressions in column 1 of Table 9 confirm that regions where union wages are larger have union workers working shorter hours. But this relationship is statistically significant only for women, for whom a ten percentage point larger union wage is associated with a .22 hour reduction in work, from the mean of 8.2 per day, for an elasticity of -.27. The regional share of union workers is associated with shorter hours, but not significantly so for either sex, based on Huber corrected standard errors.

The single equations logit models estimated in columns 2-5 in Table 9 are consistent, but because they are not jointly estimated, they are inefficient. Table 10 reports maximum likelihood estimates of the multinomial logistic model for this multiple choice problem. The estimates should be interpreted as the partial effects of the covariates on the logit index of the selected activity relative to nonparticipation (i.e. the "activity" not chosen). The estimates should be more efficient, conditional on the full information model being valid (Madalla, 1983). Conversely, misspecification in any part of the system could distort estimates throughout (Fisher, 1966).

According to the single equation logit estimates of employment in column 2 of Table 9, an increase in union workers wages of ten percentage points in a region is associated with a reduction in young female employment in that region that tends to be 4.8 percent lower, compared with a mean employment rate of 13.7 percent. Thus, for younger women the point estimate of the elasticity of employment with respect to the union relative wage effect is -.05, but it is not statistically significantly different from zero. For men the effect of a ten percentage point increase in the union relative wage effect is to decrease employment by 5.6 percent from a mean of 24.8 percent, or the elasticity of young male employment with respect to the union relative wage is -.06 and is significant at the 10 percent level. Among older workers, age 30 to 65, the union wage effect is more significant for women, as reported in Table A-3. The number of young persons unemployed decreases with the union relative wage effect for men and women, contrary to the predictions of Harris-Todaro-Calvo, and is statistically significant at the 10 percent level. The union relative wage effect is associated with increased male school enrollment, and the effect is statistically significant.

An increase in the union share of employees in a region is associated with a positive employment gain for men, but losses for women. However, this union share of jobs in the region is not significantly related to unemployment for men or women. The union share reduces the percentage of males in school, whereas it increases insignificantly females in school. Recall that a smaller proportion of African women employees are union members compared with men, 22.2 versus 37 percent (Table 1). In sum, an increase in the union share of a region's labor market shifts men's time allocation from school to employment, but it contributes to the opposite pattern for women's time allocation. The regional level of wages is included in the model in an effort to distinguish only the effect of unions and not the mix of industries or educational composition of the regional workforce. The regional wage level does not account for a significant amount of the variation in any of the categories of activity for either young men or women, though it may contribute to an increase in non participation among older women as shown in Table A-3.

Table 10 confirms similar patterns of response when a multinomial logistic model is fit to the four-way time allocation data for African men and women age 16-29. In regions where union wages are exceptionally high compared to nonunion wages, African young men and women are more likely to have left the labor force, both from employment and from unemployment. Some men tend to return to school relative to nonparticipation, whereas for women those who cease to be in the labor force apparently return to home production and leisure activities and fewer continue in school. The schooling-nonparticipation odds ratio in Table 10 is not significantly associated with the relative union wage effect. In regions where the union share of employment is larger, more men are both employed and unemployed and fewer women are in both of these labor force categories relative to nonparticipation. The same pattern is observed for older workers using the multinomial logistic model (not reported). Schooling relative to nonparticipation is positively related to the union share of employees in a region. Although the regional wage level was not associated significantly with any of the single equation logit time allocations in Table 9, it is negatively associated with male employment and unemployment relative to nonparticipation in the multinomial logit estimates in Table 10.

If the union relative wage effect across all industries for African males were somehow cut in half, from .468 (Table 3) to .234 for the mean, the single equation logistic model indicates employment for African males age 16-29 would increase by 3.3 percentage points (.234*.696), and unemployed male African youth would decline by 2.3 percentage points. The point estimates of the wage effects on African women's employment are of a similar magnitude, but are not statistically significant (Table 9), although female union workers now work noticeably longer hours at their lower wages. There does not appear to be a close local labor market connection between the union share of employment and the number of unemployed men or women, but the relative wage effect of unions is associated with lower unemployment. This observed pattern is the opposite of that predicted by the Harris-Todaro-Calvo models of labor market sectoral distortions caused by unions or government wage regulations. Thus, a reduction in the relatively higher wage received by union members is associated here with an increase in labor market participation of youth and an increase in unemployment. A reduction in male school enrollment rates is also be expected as the prospects for employment improve according to our estimates.

7. CONCLUSION

This paper has examined union wage effects in South Africa among Africans and whites, controlling for human capital variables, rural residence, and industry. The wage function estimates show that union membership among African workers increases their wages by 145 percent (exp. (.895) -1.0) at the bottom tenth percentile of the wage distribution and increases their wages by 19 percent at the top 90th percentile (Table 3). Among white workers, the relative increase in union wages is 21 percent at the tenth percentile but is associated at the 90th percentile with a reduction of 24 percent. Wage structure estimates by quantile regression also suggest that unions in South Africa are associated with lower wage returns on unobservable "ability," as well as lower returns to observed years of education and postschooling experience (Table 7). In the case of African workers, this effect comes from relatively higher union wage benefits at the lower tail of the wage residual distribution. Among white workers, the upper 70 percent of the residual distribution of wages receive lower wages in union jobs than for observationally comparable nonunion workers; there may be compensating unobserved union job amenities that white workers are willing to pay for, such as greater job security. One could also explain this pattern if the nonunion white workers were employed in what Pencavel (1995) called the administered wage sector where union wage premia are paid to retain skilled workers.

From Table 2 it can be seen that four-fifths of union members are nonwhite. It is also clear from the quantile regressions that the wage benefits from union membership are disproportionately skewed toward lower wage earners, within schooling, experience, race, and sex groups of workers. Moreover, the union relative wage gains are proportionately larger for Africans than for whites, reducing thereby the interracial disparities in wages in the union sector. Should unions in the context of South Africa then be viewed as an egalitarian force?

Seven-tenths of African wage earners are not in a union and stand to lose relatively and perhaps absolutely because of union relative wage effects. Moreover, African union workers are only one seventh of the African labor force, and only six percent of the African population age 16 to 65 (Table 1). Estimates of union relative wage differentials differ widely across countries and time periods. Lewis (1986) has estimated union wages exceed nonunion wages in the United States by about 15 percent in the 1960s and 1970s and Blanchflower (1997) does not find it has changed substantially through 1995 or differs much from the 13 percent union wage effect estimated for the U.K. by 1992-94. Conceivably, the union wage differential in South Africa might decline gradually to half its current level as the economy lowers its global protection. The hypothesized remaining 23 percent union wage effect would still be one half larger than in the U.S. or U.K. According to the estimates in Table 9, such a reduction would be associated with 1.5 percent more women age 16-29 finding employment compared with the current 14 percent and 2.3 percent more men employed compared with current 25 percent. In the age group 30-65, employment might also be expected to increase for African women (Table A-3). In sum, reducing the union relative wage effect by half could increase African employment by about two percent, roughly equal to an expansion of one-eighth in African youth employment. There would be a redistribution of wage payments from the upper middle-class African union workers to lower wage nonunion workers and the marginalized poor who are often now not actively participating in the labor force. A complete assessment of the union wage differential would require knowledge of the marginal productivity of persons displaced from the labor force to work only in home production. The effects of unions on employment opportunities for youth, and on wage returns to schooling if they find employment, may also change the gender mix of students in the schools and even the overall level of enrollments in South Africa. Further research should be able to develop more meaningful boundaries to regional labor markets as they are evolving in the new South Africa, to improve our information about the enforcement powers and reach of Industrial Councils, to incorporate industry-specific measures of international trade distortions that may help to explain the industrial dispersion in union wage effects, and finally to characterize how firm size and ownership structure affect wages. Opening up these additional avenues to research on unions and other labor market institutions in South Africa should proceed rapidly now that nationally based representative surveys for the entire country are becoming available to the public.

Notes

1. In addition, South Africa has adopted the European model of national wage-bargaining at the level of industry, and established industrial councils with the responsibility to enforce in the courts minimum wage floors on all firms in an industry, regardless of the union status of its workforce or number of employees in the firm (Moll, 1993a, 1995; Bendix, 1995). It remains unclear how effective industrial councils are in enforcing wage floors, but if they raised wages within an industry among nonunions workers, this effect would lead to an understatement of the relative unions wage effect as estimated in this paper. Our measures of industry, however, are highly aggregated into nine groupings but do reflect in Table 1 some of the industry differences in unions shares of workers as reported from incomplete data by COSATU (Baskin, 1994).

2. Specifically, we do not know what characteristic of workers would allow us to predict who is a union member, where this characteristics would not be a plausible factor also affecting the worker's productivity and hence wage offer. Such an exclusion restriction would be required to implement a sample selection bias correction of wages in either the union or nonunion subsamples, or to condition wages on endogenous union status. Identification on the basis of functional form is a debatable procedure (Heckman, 1979; Manski, 1995).

3. Transvaal, Cape, Natal, Orange, Kwazulu, Kangwane, Qwaqwa, Gazankulu, Lebowa, Kwandebele, Transkei, Bophuthatswana, Venda, and Ciskei.

4. There are no data in the survey on working conditions. Although hourly wage rates include bonuses and a variety of in-kind sources of income received by the worker, fringe benefits are not explicitly identified.

5. From means reported in Tables 9 and A-5 it can be seen that 25 percent of African men age 16-29 who are in the labor force are unemployed and 31 percent of African women.

6. This interpretation of our findings was suggested by John Pencavel.

7. Because there are no union workers in our sample of young African men in three regions (Kangwane, Kwandebele, and Natal) and for young African women in two regions (Qwaqwa and Ciskei), the union wage effect cannot be estimated, and dummies for these regions are included in the logit and OLS estimates reported in Tables 9 and 10 to retain the full sample. An alternative approach of using the older sample age 30 to 65 to estimate these regional union wage effects led to similar conclusions.

References

Baskin, J., ed., 1994. Unions in Transition: Cosatu at the Dawn of Democracy. National Labour and Economic Development Institute, Johannesburg.

Bendix, S., 1985. Industrial Relations in South Africa. Cape Town: Juta & Co. Ltd.

Blanchflower, D.G., 1997. "Changes Over Time in Union Relative Wage Effects in Great Britain and the United States," Working Paper No. 6100, Cambridge, MA: National Bureau of Economic Research.

Boal, W. and J. Pencavel, 1994. "The Effect of Labor Unions on Employment, Wages, and Days of Operation," Quarterly Journal of Economics, Vol. 109, No. 1, pp. 267-98.

Buchinsky, M., 1994. "Changes in the U.S. Wage Structure 1963-1987: Application of Quantile Regression," Econometrica, Vol. 62, No. 2, pp. 405-58.

Calvo, G., 1978. "Urban Unemployment and Wage Determination in LCDs: Trade Unions in the Harris-Todaro Model," International Economic Review, Vol. 19, No. 1, pp. 65-81.

Card, D., 1992. "The Effects of Unions on the Distribution of Wages," Working Paper No. 4195, Cambridge, MA: National Bureau of Economic Research.

Chamberlain, G., 1994. "Quantile Regression, Censoring and the Structure of Wages." In C. Sims and J.J. Laffont, eds., Proceedings of the Sixth World Congress of Econometrics Society (Barcelona, Spain). New York, NY: Cambridge University Press.

Deaton, A., 1997. The Analysis of Household Surveys, Baltimore, MD: Johns Hopkins University Press, pp. 80-85.

Duncan, G., and F. Stafford, 1980. "Do Union Members Receive Compensating Wage Differentials," American Economic Review, Vol. 70, No. 3, pp. 355-71.

Efron, B., 1979. "Bootstrap Methods: Another Look at the Jackknife," The Annals of Statistics, Vol. 7, pp. 1-26.

Farber, H.S., 1983. "The Determination of Union Status of Workers," Econometrica, Vol. 51, No. 5, pp. 1417-38.

Fisher, F.M., 1966. The Identification Problem in Econometrics. New York, NY: McGraw Hill.

Freeman, R., 1980. "Unionism and the Dispersion of Wages," Industrial and Labor Relations Review, Vol. 34, pp. 3-23.

Harris, J., and M. Todaro, 1970. "Migration, Unemployment and Development," American Economic Review, Vol. 60, No. 1, pp. 126-42.

Heckman, J., 1979. "Sample Selection Bias as a Specification Error," Econometrica, Vol. 47, No. 1, pp. 153-61.

Huber, P.J., 1967. "The Behavior of Maximum Likelihood Estimates under Non-Standard Assumptions." Proceedings of the Fifth Berkeley Symposium on Mathematical Statistics and Probability, Vol. 1, No. 1, pp. 221-33.

Hydak, T., 1979. "The Effects of Unions on Earnings Inequality in Local Labor Markets," Industrial and Labor Relations Review, Vol. 33, No. 1, pp. 77-84.

Iyengar, M., and R.C. Porter, 1990. "South Africa Without Apartheid: Estimates from General Equilibrium Simulations," Journal of International Development, Vol. 2, No. 1, pp. 1-59.

Johnson, H., and P. Mieszkowski, 1970. "The Effects of Unionisation on the Distribution of Income," Quarterly Journal of Economics, Vol. 84, pp. 539-61.

Juhn, C., K. Murphy, and B. Pierce, 1993. "Wage Inequality and the Rise in Returns to Skills," Journal of Political Economy, Vol. 101, No. 3, pp. 410-42.

Knight, J.B., 1997. "Labour Market Issues in Zimbabwe: Lessons for South Africa," The South African Journal of Economics, Vol. 65, No. 1, pp. 69-98.

Knight, J.B., and M.D. McGrath, 1977. "An Analysis of Racial Wage Discrimination in South Africa," Oxford Bulletin of Economics and Statistics, November.

Koenker, R., and G. Bassett, 1978. "Regression Quantiles," Econometric, Vol. 46, pp. 33-50.

Lemieux, T., 1993. "Unions and Wage Inequality in Canada and the United States." In D. Card and R. Freeman, eds., Small Differences that Matter. Chicago, IL: University of Chicago Press.

Lewis, H.G., 1963. Unionism and Relative Wages in the United States. Chicago, IL: University of Chicago Press.

_____, 1986. Union Relative Wage Effects: A Survey. Chicago, IL: University of Chicago Press.

Lipton, M., 1985. Capitalism and Apartheid: South Africa, 1910-1984. Aldershot: Gower Publ. Co.

Maddala, G.S., 1983. Limited-Dependent and Qualitative Variables in Econometrics. Cambridge: At the University Press.

Manski, C.F., 1995. Identification Problems in the Social Sciences. Cambridge, MA: Harvard University Press.

Mincer, J., 1974. Schooling, Experience and Earnings. New York, NY: Columbia University Press.

_____, 1976. "Unemployment Effects of Minimum Wages," Journal of Political Economy, Vol. 84, No. 4, pt. 2, pp. S87-S104.

Moll, P.G., 1993a. "Black South African Unions: Relative Wage Effects in International Perspective," Industrial and Labor Relations Review, Vol. 46, No. 2, pp. 245-61.

_____, 1993b. "Industry Wage Differentials and Efficiency Wages," Journal of Development Economics, Vol. 41, pp. 213-46.

_____, 1996. "Compulsory Centralization of Collective Bargaining in South Africa," American Economic Review, Vol. 86, No. 2, pp. 326-29.

Mwabu, G., and T.P. Schultz, 1995. "Wage Premia for Education and Location, by Gender and Race in South Africa," Unpublished paper, Economic Growth Center, Yale University.

_____, 1996. "Education Returns Across Quantiles of the Wage Function," American Economic Review, Vol. 86, No. 2, pp. 335-39.

Pencavel, J., 1991. Labor Markets under Trade Unionism. Cambridge, MA: Blackwell Pub.

_____, 1995. "The Legal Framework for Collective Bargaining in Developing Countries," Unpublished paper, Stanford University.

Porter, R., 1976. "A Model of the Southern African-Type Economy," American Economic Review, Vol. 68, No. 5, pp. 734-55.

SALDRU, 1994. South Africans Rich and the Poor: Baseline Household Statistics. Cape Town: University of Cape Town, School of Economics.

Schultz, T.P., 1980. "Estimating Labor Supply Functions for Married Women." In J.P. Smith, ed., Female Labor Supply. Princeton, NJ: Princeton University Press.

_____, 1982. "Effective Protection and the Distribution of Personal Income by Sector in Colombia." In A. Krueger, ed., Trade and Employment in Developing Countries. Chicago, IL: University of Chicago Press.

Standing, G., J. Sender, and J. Weeks, 1996. Restructuring the Labour Market: The South Africa Challenge. Geneva, Switzerland: International Labor Office.

Thompson, L., 1990. A History of South Africa. New Haven, CT: Yale University Press.

White, H., 1980. "A Heteroscedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroscedasticity," Econometrica, Vol. 48, No. 4, pp. 817-38.

Table 1

Number of Employed Respondents in 1993 LSMS by Industry, Race and Sex, with Percent Union Member (in parentheses)

Industrial Sector	African		Colored		Asian		White		All Races Both Sexes
	Male	Female	Male	Female	Male	Female	Male	Female	
1. Agriculture and Fishery	462 (2.8)	175 (3.4)	48 (16.7)	34 (14.7)	-	-	28 (10.7)	7 (0)	754 (4.5)
2. Mining	395 (76.2)	7 (100.)	8 (100.)	-	2 (50.)	-	68 (55.9)	10 (30.)	490 (73.1)
3. Manufacturing	41.7 (54.7)	198 (41.9)	91 (55.0)	74 (62.0)	56 (64.3)	45 (55.6)	106 (18.9)	36 (8.3)	1023 (48.0)
4. Electricity and Water	60 (28.3)	1 (0)	12 (66.7)	-	2 (0)	1 (0)	35 (48.6)	6 (16.7)	117 (36.8)
5. Construction	176 (25.6)	16 (12.5)	34 (20.6)	3 (100.)	12 (33.3)	1 (0)	60 (18.3)	5 (0)	307 (23.5)
6. Wholesale/Retail Restaurant/Hotel	249 (30.9)	204 (32.8)	36 (27.8)	67 (38.8)	30 (16.7)	16 (37.5)	60 (6.7)	93 (4.3)	755 (26.4)
7. Transportation, Communication, and Finance	196 (39.3)	44 (34.1)	44 (38.6)	15 (66.7)	20 (40.0)	9 (11.1)	173 (22.0)	119 (16.0)	620 (29.8)
8. Educational, Medical, and Legal Services	210 (33.3)	385 (35.1)	28 (60.7)	70 (48.6)	23 (43.5)	29 (24.1)	48 (4.2)	162 (9.3)	955 (30.4)
9. Domestic and Other Services and Other Industry	199 (20.9)	503 (4.4)	49 (34.7)	49 (16.3)	25 (36.0)	6 (33.3)	75 (10.7)	89 (13.5)	955 (12.0)
All Industries	2364 (36.8)	1533 (22.0)	350 (40.6)	312 (42.3)	170 (42.9)	107 (38.3)	653 (21.6)	527 (10.8)	6016 (29.8)

Table 2

Characteristics of the South African 1993 Survey Population Age 16-65
and of the Wage Earner Samples, by Race and Sex

Sample Size Variable Mean	African		Colored		Indian		White	
	Men	Women	Men	Women	Men	Women	Men	Women
All Persons Age 16-65:	9325	10473	958	1075	363	411	1447	1517
Years of Education	7.00	6.92	8.62	8.17	10.6	9.34	11.8	10.4
Proportion of Population with Some Education:								
Primary	.874	.843	.959	.944	.961	.927	.955	.916
Secondary	.493	.506	.720	.644	.909	.796	.928	.840
Higher	.023	.026	.031	.026	.140	.061	.346	.197
Post-School Experience in Years	19.5	20.2	18.5	19.5	18.3	18.8	18.9	20.6
Rural Resident	.676	.676	.072	.063	.008	.005	.092	.090
Homelands	.640	.687	.001	.001	.006	.005	.001	.001
Proportion of Population in Labor Force	.521	.318	.724	.558	.786	.392	.850	.622
Proportion of Labor Force Unemployed	.145	.175	.123	.180	.072	.090	.028	.039
Proportion of Population Wage Earners	.254	.146	.365	.290	.468	.260	.451	.347
Proportion of Population in School	.193	.177	.133	.117	.135	.129	.129	.105
Wage Earner Sample Size:	2364	1533	350	312	170	107	653	527
Years of Education	6.81	7.55	8.54	8.51	11.0	10.7	12.2	11.2
Post-Schooling Experience in Years	24.7	24.2	21.2	20.3	18.9	16.2	19.2	18.6
Rural Resident	.561	.472	.103	.099	.018	.009	.066	.068
Homelands	.335	.419	0	0	.012	.009	0	.002
Proportion of Union Members	.368	.220	.406	.423	.429	.383	.216	.108
Log Hourly Wage Rate	1.56	1.18	1.97	1.62	2.55	2.10	3.33	2.76
Hourly Wage in Rands ($1.00 = 4.64 Rand; Financial Rate 5/93)	4.76	3.25	7.17	5.05	12.81	8.17	27.9	15.80

Source: Authors calculation from the survey file for all persons age 16-65 reporting their sex, education, residential region and wage earner identifier variables.

Table 3

Quantile Regression Estimates of the Wage Function
Controlling for Union Membership for African Males
(absolute values of Bootstrap t-ratios in parentheses)

Explanatory Variables	Wage Quantiles .10	.50	.90	Mean (OLS)
A. Education and Experience Variables				
Years of Primary Education	.085 (5.56)	.075 (7.59)	.033 (3.14)	.074 (9.74)
Years of Secondary Education	.143 (6.67)	.142 (11.2)	.169 (13.6)	.160 (14.9)
Years of Higher Education	.300 (7.14)	.322 (8.45)	.274 (6.56)	.293 (10.2)
Potential Job Experience in Years	.049 (3.68)	.053 (8.25)	.049 (4.20)	.049 (9.27)
Potential Job Experience Squared ($x10^{-2}$)	-.0720 (3.73)	-.0687 (5.67)	-.0638 (3.34)	-.0655 (7.19)
B. Location and Union Variables				
Rural Area (1=Rural Residence)	-.437 (5.62)	-.298 (9.00)	-.237 (9.20)	-.357 (11.1)
Union Status (1=Union Member)	.895 (13.9)	.446 (13.1)	.107 (1.84)	.468 (14.7)
Constant Term	-.800 (.418)	.180 (2.05)	1.291 (7.58)	.219 (2.41)
Pseudo R^2	.321	.217	.215	.399
Sample Size	2364			

Table 4

Quantile Regression Estimates of the Wage Function
Controlling for Union Membership for White Males
(absolute values of Bootstrap t-ratios in parentheses)

Explanatory Variables	Wage Quantiles			
	.10	.50	.90	Mean (OLS)
A. Education and Experience Variables				
Years of Primary Education	.016 (.06)	-.056 (1.82)	-.036 (.52)	-.011 (.45)
Years of Secondary Education	.242 (4.18)	.090 (1.92)	.007 (.25)	.084 (3.01)
Years of Higher Education	.098 (3.56)	.126 (6.15)	.183 (6.23)	.150 (8.34)
Potential Job Experience in Years	.126 (6.61)	.078 (6.74)	.091 (4.91)	.103 (10.9)
Potential Job Experience Squared ($x10^{-2}$)	-.250 (5.15)	-.130 (4.71)	-.149 (3.80)	-.187 (8.89)
B. Location and Union Variables				
Rural Area (1=Rural Residence)	-.057 (.06)	-.178 (1.81)	-.361 (2.41)	-.294 (2.62)
Union Status (1=Union Member)	.188 (1.88)	-.112 (2.33)	-.270 (3.32)	-.051 (.75)
Constant Term	.104 (.06)	2.344 (16.0)	3.032 (5.84)	1.823 (11.2)
Pseudo R^2	.261	.171	.160	.276
Sample Size	653			

Table 5

Quantile Regression Estimates of the Wage Function Controlling for
Union Membership and Employment Industry for African Males
(absolute values of Bootstrap t-ratios in parentheses)

Explanatory Variables	Wage Quantiles			
	.10	.50	.90	Mean (OLS)
A. Education and Experience Variables				
Years of Primary Education	.024 (2.93)	.055 (6.77)	.026 (2.89)	.043 (6.17)
Years of Secondary Education	.126 (6.20)	.113 (6.36)	.119 (6.19)	.118 (11.6)
Years of Higher Education	.259 (7.27)	.272 (6.59)	.273 (4.93)	.255 (9.60)
Potential Job Experience in Years	.056 (4.38)	.044 (6.93)	.040 (4.22)	.044 (9.31)
Potential Experience Squared (x 10^{-2})	-.0864 (3.81)	-.0543 (4.84)	-.0560 (3.51)	-.0613 (7.53)
B. Location and Union Variables				
Rural Area (1=Rural Residence)	-.194 (3.95)	-.189 (5.38)	-.110 (2.27)	-.215 (6.85)
Union Status (1=Union Member)	.345 (4.56)	.190 (5.13)	.005 (.10)	.191 (5.89)
C. Employment Industries [Manufacturing is the excluded category]				
Agriculture, Forestry and Fisheries	-1.213 (17.3)	-1.042 (16.4)	-.566 (7.31)	-.988 (19.3)
Mining	.193 (2.22)	.069 (1.59)	-.077 (1.12)	.079 (1.64)
Construction	-.363 (2.23)	-.237 (3.25)	-.128 (1.01)	-.239 (4.01)
Wholesale and Restaurants	-.243 (3.10)	-.135 (2.71)	-.174 (2.42)	-.165 (3.11)
Transport, Communications and Finance	-.124 (1.39)	.147 (.19)	.046 (.45)	-.007 (.11)
Education, Medical and Legal Services	.222 (3.47)	.103 (1.53)	.125 (1.11)	.142 (2.35)
Domestic and Other Services	-.632 (4.87)	-.407 (4.90)	-.362 (2.89)	-.435 (6.39)
Armed Forces, Electricity, Water and Other	-.014 (.14)	.009 (.13)	.239 (2.69)	.077 (1.17)
Constant Term	.020 (.11)	.673 (6.27)	1.658 (10.4)	.780 (8.74)
Pseudo R^2	.441	.313	.262	.524
Sample Size	2364			

Table 6

Quantile Regression Estimates of the Wage Function Controlling for Union Membership and Employment Industry for White Males
(absolute values of Bootstrap t-ratios in parentheses)

Explanatory Variables	Wage Quantiles .10	.50	.90	Mean (OLS)
A. Education and Experience Variables				
Years of Primary Education	-.011 (.06)	-.057 (2.13)	-.033 (.53)	-.012 (.49)
Years of Secondary Education	.239 (4.15)	.080 (3.38)	.005 (.18)	.082 (2.94)
Years of Higher Education	.101 (3.64)	.136 (6.74)	.155 (6.56)	.151 (8.31)
Potential Job Experience in Years	.122 (2.14)	.080 (7.80)	.086 (6.80)	.101 (10.6)
Potential Experience Squared (x10^{-2})	-.253 (4.15)	-.137 (5.80)	-.138 (5.32)	-.185 (8.70)
B. Location and Union Variables				
Rural Area (1=Rural Residence)	-.226 (2.29)	-.263 (1.60)	-.373 (1.31)	-.382 (2.96)
Union Status (1=Union Member)	.142 (1.51)	-.119 (1.86)	-.249 (2.18)	-.097 (1.35)
C. Employment Industries [Manufacturing is the excluded category]				
Agriculture, Fisheries and Forestry	-.734 (1.30)	-.205 (1.56)	.169 (.72)	-.279 (1.84)
Mining	.088 (.66)	.157 (1.28)	.113 (.50)	.168 (1.36)
Construction	-.327 (2.32)	.008 (.07)	.304 (1.59)	.002 (.02)
Wholesale and Restaurants	-.071 (.41)	-.055 (.49)	.308 (1.72)	.064 (.55)
Transport, Communications and Finance	-.074 (.71)	.003 (.03)	.088 (1.10)	-.027 (.31)
Education, Medical and Legal Services	-.182 (.95)	-.130 (.21)	.016 (.05)	-.048 (.39)
Domestic and Other Services	-.436 (2.37)	-.032 (.21)	-.089 (.54)	-.182 (1.53)
Armed Forces, Electricity, Water and Other	-.069 (.65)	.045 (.48)	-.017 (.17)	-.020 (.17)
Constant Term	.510 (.44)	2.401 (14.2)	3.011 (6.24)	1.889 (10.6)
Pseudo R^2	.281	.178	.181	.288
Sample Size	653			

Table 7

Quantile Regression Estimates of the Wage Function Controlling for Union Membership Interacted with Education and Experience Coefficients for African Males

Explanatory Variables	Wage Deciles .10	Wage Deciles .50	Wage Deciles .90	Mean (OLS)
Years of Primary Education	.0570 (2.92)	.0783 (5.12)	.0278 (2.26)	.0849 (9.56)
Years of Primary Education *Union	-.0179 (.52)	-.0643 (4.03)	-.0052 (.31)	-.0510 (3.14)
Years of Secondary Education	.261 (8.71)	.201 (11.9)	.177 (9.39)	.203 (15.4)
Years of Secondary Education *Union	-.228 (4.55)	-.133 (6.99)	-.0364 (1.47)	-.115 (5.27)
Years of Higher Education	.213 (4.43)	.305 (8.00)	.264 (5.03)	.296 (8.50)
Years of Higher Education *Union	.0356 (.51)	-.0938 (1.63)	-.182 (3.64)	-.0240 (.41)
Potential Job Experience in Years	.0669 (4.74)	.0602 (5.12)	.0470 (3.65)	.0542 (8.53)
Job Experience *Union	-.0279 (.94)	-.0451 (2.94)	-.0172 (.90)	-.0189 (1.70)
Potential Job Experience Squared ($x10^{-2}$)	-.103 (3.79)	-.0829 (4.30)	-.0694 (3.24)	-.0737 (6.84)
Job Experience Squared *Union ($x10^{-2}$)	.0384 (.70)	.0680 (2.77)	.0309 (.24)	.0299 (1.54)
Rural Area (=1)	-.535 (5.83)	-.553 (12.5)	-.359 (7.76)	-.352 (11.1)
Union Status (1=Union Member)	1.70 (4.38)	1.68 (7.00)	.356 (1.17)	1.16 (6.64)
Constant Term	-1.00 (5.34)	-.0540 (.32)	1.39 (6.69)	.0449 (.43)
Pseudo R^2	.361	.254	.232	.420

Table 8

Quantile Estimates of the Wage Function with Union Status interacted with Industry As Well As Education and Experience Variables for African Males

Explanatory Variables[a] (Mining Industry Omitted)	Wage Deciles			Mean OLS
	.10	.50	.90	
Agriculture	-1.67 (26.4)	-1.38 (16.4)	-.566 (6.96)	-1.35 (18.2)
Agriculture *Union	.930 (6.33)	.571 (1.79)	.122 (.59)	.889 (4.55)
Manufacturing	-.529 (3.74)	-.630 (5.44)	-.0681 (.52)	-.340 (4.10)
Manufacturing *Union	.503 (3.19)	.560 (5.10)	.123 (.83)	.345 (3.45)
Construction	-.687 (2.73)	-.725 (5.26)	-.201 (.82)	-.629 (7.12)
Construction *Union	.636 (2.33)	.808 (3.12)	.883 (2.13)	.528 (3.90)
Wholesale/Retail Trade Hotels	-.849 (8.69)	-.811 (8.48)	-.341 (2.44)	-.552 (6.54)
Trade *Union	.636 (2.91)	.600 (5.89)	.399 (2.10)	.500 (4.36)
Transportation, Communication and Finance	-.817 (7.79)	-.804 (7.63)	-.233 (3.57)	-.308 (3.38)
Transport *Union	.729 (4.77)	.748 (7.84)	.216 (1.72)	.255 (2.09)
Education, Medical and Legal Services	-.624 (6.23)	-.474 (4.97)	-.0217 (.13)	-.202 (2.26)
Education, etc. *Union	.371 (2.59)	.416 (3.51)	.197 (1.31)	.380 (3.09)
Domestic Services	-.350 (2.65)	-.401 (3.34)	-.149 (.90)	-1.22 (9.76)
Domestic Union	.555 (3.31)	.798 (4.36)	.318 (1.43)	.718 (2.26)
Other Services and Utilities	-1.02 (8.10)	-.795 (7.36)	-.0693 (.51)	-.381 (4.44)
Other Services *Union	.643 (2.55)	.673 (5.40)	.372 (1.69)	.390 (3.05)
Union-Member	.112 (1.24)	-.0266 (.27)	-.0251 (.23)	-.153 (2.00)
Rural Residence	-.106 (2.65)	-.343 (7.02)	-.126 (1.77)	-.222 (7.13)
Constant	.396 (3.65)	1.19 (8.21)	1.66 (11.6)	.588 (11.0)
Pseudo R^2	.442	.320	.266	.534

[a] Education spline and quadratic in experience are also included, and interacted with union status, but not reported to save space.

Table 9

Estimated Effects of Union Share and Union Relative Wage in the Local Labor Market on Economic Activity of Young Africans, Age 16-29 (Huber standard errors are the basis for the reported t-ratio in parentheses)

Selected Explanatory Variables[a]	Male					Female				
	Hours Worked Per Day[d] Regression (1)	Employed in Survey Week Logit (2)	Unemployed Logit (3)	Enrolled in School Logit (4)	Not in Labor Force nor in School Logit (5)	Hours Worked per Day[d] Regression (1)	Employed in Survey Week Logit (2)	Unemployed Logit (3)	Enrolled in School Logit (4)	Not in Labor Force nor in School Logit (5)
Union Effect on Log Wage by Region[b]	-.789[d] (.97)	-.696 (1.65)	-1.55 (3.83)	.985 (7.51)	.314 (1.23)	-2.22[d] (5.92)	-.646 (.58)	-2.99 (1.97)	-.022 (.11)	.772 (1.29)
Share of Wage Employees who are Union Members by Region	-1.90 (1.57)	6.46 (5.29)	-1.36 (1.07)	-2.36 (2.40)	-4.23 (6.59)	-1.09 (1.02)	-5.54 (2.20)	-1.66 (.64)	2.14 (2.92)	1.34 (.96)
Regional Level in Log Wage[c]	.623 (.48)	-.498 (.39)	-1.28 (1.11)	-.007 (.01)	.779 (1.41)	-.113 (.21)	-.232 (.44)	.151 (.23)	.067 (.18)	.087 (.23)
Pseudo R^2	.050	.249	.085	.470	.076	.050	.139	.073	.433	.137
Sample Size	749	3848	3848	3848	3848	439	4533	4533	.4533	4533
Mean of Dependent Variable	9.00	.248	.082	.430	.240	8.23	.137	.062	.389	.412

[a] Controls also included for education, age, age squared, Qwaqwa and Ciskei province for females, and Kangwane, Kwandebele, and Natal for males, because there were no African union employed workers age 16-29 in these regions on which to estimate the variable "union effect on wages."
[b] The union effect on wages is the estimated coefficient for the union members regional dummy variable in a log wage equation that is also conditioned on years of education, age, age squared, ten rural-regions interactions, and industry dummies. The variable estimates the deviation of a province's log wage of union from the log wage of a nonunion worker.
[c] Provincial wage effects are designed to approximate regional labor market levels of wages by the value of provincial dummy coefficients in a log wage regression, controlling for education and a quadratic in age, and industry dummies.
[d] In the hours per day worked regression, the coefficient on the regional union wage effect is that on the interaction of the regional union wage effect multiplied by one if the worker is a union member and by zero otherwise. Union member status is also controlled in the hours regression (but not reported because insignificant).

Table 10

Multinomial Logistic Estimates of Economic Activity Responses of Africans Age 16 to 29 to Local Union and Labor Market Variables[a]

Selected Explanatory Variables	African Males (n = 3848)			African Females (n = 4533)		
	Employment/ Not Active	Unemployment/ Not Active	School/ Not Active	Employment/ Not Active	Unemployed/ Not Active	School/ Not Active
Union Wage Effect in Region	-.869 (3.55)	-1.72 (5.04)	.307 (.40)	-1.04 (3.02)	-3.29 (5.86)	-.349 (1.34)
Union Share in Region	7.15 (10.6)	2.19 (2.37)	.531 (2.28)	-5.59 (6.10)	-2.54 (1.89)	1.30 (1.91)
Regional Wage Level	-.930 (2.08)	-1.60 (2.74)	-.359 (.83)	-.187 (.69)	.146 (.36)	.0314 (.13)
Log Likelihood	-3431.98			-3923.42		
χ^2 (27 and 24 df) (p value)	2803. (.000)			2827. (.000)		
Pseudo R^2	.290			.265		

[a] Controls as in Table 9 also included for quadratic in age, years of completed education, and necessary province dummies.

APPENDIX: QUANTILE WAGE REGRESSION AND BOOTSTRAP STANDARD ERRORS

In a simple parametric regression model of (log) wage earnings (e.g., Mincer, 1974), the parameter on years of schooling is the "return" that minimizes the sum of squared deviations of sample (log) wages from their mean, conditional on sample values of schooling and other covariates. A major drawback to the least squares estimator is that it is sensitive to distribution of wages from their conditional means. The maximum likelihood estimator has the additional limitation that it assumes normality of the distribution of wage deviations, an assumption that might not hold in the sample.

We estimate wage returns to schooling by a quantile technique which minimizes the sum of absolute deviations of sample wages from a given wage quantile. This quantile regression technique has the advantage that it is robust to outliers in wages. It is also flexible in that it does not impose a particular functional form on the residuals to the wage equation. The estimated coefficients are the parameters of an unknown nonlinear wage function that best approximates a linear function.

Adapting the estimation strategies of Armstrong, Frome and Kung (1979), Barrodale and Roberts (1973), Buchinsky (1994), Chamberlain (1994), and Wagner (1959) the problem of minimization of the sum of absolute deviations of sample wages from an arbitrarily chosen quantile wage can, in the simplest case, be expressed as:

$$\text{Minimize } \Sigma_i \left| Y_{\tau i} - \Sigma_j \beta_{\tau j} X_{ij} \right|, \tag{1}$$

where,

$Y_{\tau i}$ = Wage of individual i at quantile τ, $0<\tau<1$, and $i=1,...,n$;

X_{ij} = Covariate j (e.g., schooling) for individual i, j=1,...,M;

$\beta_{\tau j}$ = Effect of covariate j on the wage rate at quantile τ.

In expression (1), the wage rate, Y, is assumed to be a linear function of βs, but the expression also holds when Y is any nonlinear function of the same parameters. The estimation problem in (1) is to find values of βs that minimize the sum of absolute deviations of wages at a given quantile, conditional on sample values of the covariates. The estimation is implemented by treating expression (1) as a linear programming problem and rewriting it as:

$$\text{Minimize } \Sigma_i(P_i + N_i), \tag{2}$$

subject to,

$$Y_{\tau i} = \Sigma_j \beta_{\tau j} X_{ij} + P_i - N_i, \tag{3}$$

and $P_i \geq 0$, $N_i \geq 0$, i,=1,...,n,

where P_i and N_i are to be interpreted as vertical deviations above and below the fitted line, respectively, so that $(P_i + N_i)$ is the absolute deviation between the fitted line, $\Sigma_j \beta_{\tau j} X_{ij}$, and the sample line, $Y_{\tau i}$. Expression (2) can be written in matrix-vector notation (see Chamberlain, 1994, p. 13) as:

$$\text{Minimize } 1/n\Sigma_i[\tau\ell(y_i \geq \beta'x_i) + (1-\tau)\ell(y_i < \beta'x_i)]\,|y_i - \beta'x_i|, \tag{4}$$

where $\ell(.)$ is an indicator function that equals one if event (.) is true and equals zero otherwise, and n is sample size. If, for example, $y_i \geq \beta'x_i$, the deviation is "positive" (at least above the fitted line) and is weighted by τ; similarly, if $y_i < \beta'x_i$, the deviation is "negative" (below the fitted line) and is weighted by $1-\tau$, so that the quantiles other than the median are estimated by weighting the regression residuals, the weight on positive and negative residuals depending on their location relative to the median residual. In finding values of βs via the linear programming technique, the above deviations are treated as nonnegative numbers as in restriction (3) because the program seeks to minimize the sum of their absolute values (see, STATA, 1995).

The standard errors of the estimated quantile regression coefficients are typically computed by the method of Koenker and Bassett (1982). These standard errors, however, are downward biased because they do not take into account the heteroscedasticity of the disturbance terms. We follow instead a bootstrap approach to estimation of the standard errors that selects bootstrap samples from the original sample with replacement (Chamberlain, 1994; Buchinsky, 1994). For each bootstrap sample, the linear programming algorithm estimates regression coefficients at each quantile. The means of the estimated bootstrap coefficients are then used to calculate their variances, $V(\beta_\tau)$, and the associated standard errors (see Efron and Tibshirani, 1993; Chamberlain, 1994) using the expression:

$$V(\beta_\tau) = n/B\Sigma_b(\beta_\tau^{(b)} - \beta_\tau)(\beta_\tau^{(b)} - \beta_\tau)' \qquad (5)$$

where B = Number of bootstrap replication samples, b=1,...,B;

$\beta_\tau^{(b)}$ = parameter estimate from bootstrap replication b;

β_τ = mean of all parameters obtained from bootstrap replication samples.

APPENDIX ADDITIONAL REFERENCES

Armstrong, Ronald D., E. I. Frome, and D. S. Kung, 1979. "A Revised Simplex Algorithm for the Absolute Deviations Curve Fitting Problem," Communications in Statistics, Simulation and Computation, Vol. B8, No. 2, pp. 175-90.

Barrodale, I., and F. D. K. Roberts, 1973. "An Improved Algorithm for Discrete l_1 Linear Approximation," SIAM Journal of Numerical Analysis, Vol. 10, No. 5, pp. 839-48.

Efron, B., and R. J. Tibshirani, 1993. An Introduction to the Bootstrap, New York, NY: Chapman and Hall.

Koenker, R., and G. Basset, 1982. "Robust Tests for Heteroscedasticity Based on Regression Quantiles," Econometrica, Vol. 50, pp. 43-61.

STATA, 1995, Reference Manual (Stata Release 4.0), Stata Corporation, University Drive, East College Station, Texas.

Wagner, Harvey M., 1959, "Linear Programming Techniques for Regression Analysis," Journal of American Statistical Association, Vol. 54, March, pp. 206-12.

Table A-1

Number of Employed Respondents in 1993 SMLS by Industry and Employment for African and White Males, with Percent Union Members (in parentheses)

Sample Group and Employment Group	Agriculture and Fishing	Mining	Manufacturing	Construction	Sales and Trade	Transport Communication Finance, and Utilities	Professional Services	Personal Services and Other
African Males:								
Private Corporation	433 (2.3)	369 (77.0)	377 (56.2)	132 (27.3)	210 (28.6)	131 (38.2)	24 (20.8)	56 (26.8)
Public Corporation	3 (0)	18 (61.1)	30 (46.7)	20 (25.)	29 (58.6)	62 (43.6)	2 (0)	9 (11.1)
Government (3 levels)	17 (17.7)	17 (17.6)	7 (71.4)	20 (20.)	6 (0)	57 (29.8)	180 (35.6)	102 (22.5)
Household Employees	5 (0)	0	0	0	0	1 (0)	0	30 (6.7)
Self Employed	2 (0)	1 (0)	0	3 (0)	3 (0)	1 (0)	1 (0)	0
Non-Profit and Other	3 (0)	1 (0)	0	1 (0)	1 (0)	1 (0)	3 (33.3)	1 (0)
White Men:								
Private Corporation	17 (5.9)	55 (58.2)	93 (18.3)	42 (19.0)	51 (7.84)	150 (22.0)	8 (0)	28 (10.7)
Public Corporation	2 (0)	6 (50.)	4 (.25)	25 (.24)	2 (0)	1 (0)	2 (0)	1 (0)
Government (3 levels)	6 (33.3)	6 (50.)	2 (100.)	6 (16.7)	3 (0)	30 (30.)	31 (6.5)	40 (12.5)
Household Employee	0	0	0	1 (100.)	0	0	0	2 (0)
Self Employed	1 (0)	0	0	0	0	0	2 (0)	3 (0)
Non-Profit and Other	1 (0)	0	0	1 (0)	0	0	2 (0)	3 (0)

Table A-2

Union Effects on Wage Levels and Union Share of Employment by Old Province - Regions for African Wage Earners Age 16-29

Province or Region (old boundaries)	Sample Size Age 16-29		Log Wage Deviation[a] Employees Age 16-29		Union Share of Employment	
	Male	Female	Male	Female	Male	Female
Transvaal	1153	996	0.0	0.0	.365	.206
Cape	243	264	.137 (.55)	.001 (.00)	.295	.121
Natal	118	160	—	.433 (.43)	.205	.141
Orange	340	353	.491 (2.58)	.154 (.38)	.361	.124
Kwazulu	557	952	-.068 (.36)	.152 (.58)	.318	.182
Kangwane	82	133	—	.355 (.75)	.136	.068
Qwaqwa	22	26	.103 (1.52)	—	.333	.143
Gazankulu	89	170	.098 (.19)	.110 (.17)	.245	.080
Lebowa	253	522	.739 (2.23)	.146 (.42)	.161	.090
Kwandebele	69	77	—	.258 (.31)	.244	.125
Transkei	295	592	.117 (.22)	.108 (.17)	.198	.204
Bophuthatswana	275	391	.156 (.58)	.533 (1.03)	.167	.105
Venda	58	97	.027 (.05)	.873 (1.89)	.125	.152
Ciskei	90	144	.323 (.34)	—	.173	.098
Union Member Effect (overall or transvaal)	n.a.	n.a.	-.011 (.10)	.242 (1.35)	n.a.	n.a.
Sample Size	3848	4533	639	380	2896	1955
R^2	n.a.	n.a.	.548	.663	n.a.	n.a.

[a] Controlling for three level spline in years of completed schooling, quadratic in experience, rural, and ten rural region dummies, ten-region dummies, and eight industry dummies.

n.a. not appropriate

— no union members in region.

Table A-3

Estimated Logit Effects of Union Share and Wage in the
Local Labor Market on Economic Activity of Africans Age 30-65
(Huber standard error are the basis for reported t-ratios in parentheses)

Selected Explanatory Variable	Male Sample Size = 3644			Female Sample Size = 4877		
	Employed	Unemployed	Not Participant	Employed	Unemployed	Not Participant
Union Effect on Log Wage	.105 (.21)	-.0918 (.13)	-.0988 (.19)	-1.00 (1.65)	-1.65 (2.77)	1.34 (2.45)
Union Share of Employees	4.73 (1.76)	-1.32 (.58)	-4.96 (1.98)	.122 (.04)	.301 (.14)	-.132 (.04)
Regional Level of Log Wage	-.572 (.32)	1.05 (.63)	.277 (.16)	-1.18 (1.61)	-1.05 (1.21)	1.39 (2.06)
Pseudo R^2	.072	.023	.087	.075	.066	.098

Table A-4

Proportion of Male Employees in Unions, By Race, Age and Education

Educational Attainment in Years	African Aged:		White Aged:	
	16-29	30-65	16-29	30-65
0	.10 (89)[a]	.21 (439)	.19 (16)	.17 (29)
1-3	.17 (46)	.25 (156)	0 (1)	0 (1)
4-7	.21 (312)	.31 (772)	.33 (3)	.25 (12)
8-11	.28 (322)	.34 (780)	.21 (52)	.28 (204)
12	.29 (162)	.35 (210)	.11 (108)	.15 (234)
13-15	.29 (17)	.29 (28)	.16 (77)	.20 (199)
16 or more	.40 (5)	.14 (14)	.05 (19)	.06 (139)
All Educations	.24 (953)	.30 (2399)	.15 (276)	.18 (818)

[a]Reported in parentheses beneath the proportion of employees in union is the number of survey respondents.

Table A-5

Descriptive Statistics for Cross Regional Analysis Reported in Tables 9, 10, and A-3[a]

Variable	Age 16-29		Age 30-65	
	Male	Female	Male	Female
Union Member	.058	.018	.197	.059
Employed in Survey Week	.248	.137	.658	.381
Unemployed	.082	.062	.070	.051
Enrolled in School	.430	.389	—	—
Not Participating in the above	.240	.412	.271	.569
Schooling Completed in Years	7.99 (3.89)	8.43 (3.11)	5.96 (3.95)	5.30 (3.97)
Age	21.8 (3.89)	21.9 (3.91)	43.0 (9.87)	44.0 (10.4)
Regional Variables:				
Union Effect on Log Wage of Region	.156 (.264)	.169 (.184)	.0930 (.193)	.247 (.272)
Union Share of Employees in Region	.208 (.093)	.150 (.075)	.286 (.084)	.154 (.048)
Region Predicted Wage without Union Effect	-.0962 (.158)	.0076 (.206)	.0378 (.116)	-.0361 (.174)
Sample Size	3848	4533	3644	4877

[a] In parentheses beneath the mean of continuous variables is the standard deviation. For binary variables
the standard deviation is $\sqrt{m(1-m)}$, where m is the reported mean.