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Abstract

This paper examines the link between trade protection and industry wage premia in India using a unique dataset combining employment survey data with industry-level data for various years between 1983 and 2000. The author finds that India's trade reforms were not distributionally neutral. The impact of protection on industry wage premia was positive and statistically significant, though modest in magnitude: workers employed in industries with high tariffs received higher wages than apparently identical workers in low-tariff industries. Because industries with high initial levels of protection were also those with the largest tariff reductions during this period and had the highest share of unskilled workers, the positive tariff-wage effect implies that the trade reforms were likely to have increased wage inequality as the relative wages of the (predominantly unskilled) workers in these manufacturing industries fell.

KEYWORDS: Trade Protection and Wages, India

TRADE PROTECTION AND INDUSTRY WAGES IN INDIA

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This paper examines the link between trade protection and industry wage premia in India using a unique dataset combining employment survey data with industry-level data for various years between 1983 and 2000. The author finds that India's trade reforms were not distributionally neutral. The impact of protection on industry wage premia was positive and statistically significant, though modest in magnitude: workers employed in industries with high tariffs received higher wages than apparently identical workers in low-tariff industries. Because industries with high initial levels of protection were also those with the largest tariff reductions during this period and had the highest share of unskilled workers, the positive tariff-wage effect implies that the trade reforms were likely to have increased wage inequality as the relative wages of the (predominantly unskilled) workers in these manufacturing industries fell.

The 1990s were a period of rapid trade liberalization and industrial deregulation in India. Despite considerable debate regarding these reforms, little systematic empirical work has investigated their wage effects. This paper is among the first attempts to fill this gap through an econometric examination of the link between trade protection and industry wage premia.

The modest but growing literature in this field yields ambiguous conclusions on the impact of trade protection on relative industry wages, and the primary goal of this paper is to present further empirical evidence for India. Several studies have suggested that workers' industry affiliation is an important determinant of the wage, either because of returns to industry-specific skills that cannot be transferred in the short- to medium-run or because of industry rents arising out of imperfect competition (Krueger and Summers 1988). There is some evidence that labor reallocation in the wake of trade liberalization is limited in developing countries, possibly due to labor market rigidities, suggesting that adjustment to trade reforms might occur through wage rather than employment channels for industries that experience relative tariff changes (Goldberg and Pavcnik 2004).

This paper draws on data from three large-scale employment surveys conducted in 1983, 1993–94, and 1999–2000. Industry wage premia are obtained by filtering out the

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The dataset compiled for the trade and wage premia regressions and a data appendix with additional results are available from Puja Vasudeva Dutta at the World Bank, 70 Lodi Estate, New Delhi—110003, India. Telephone: +91(11)41479156; fax: +91(11)24619393; email: pdutta@worldbank.org.

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effects of observable worker characteristics using wage regression models that control for potential selection bias. The role of trade policy in determining these estimated wage premia is then assessed. The empirical analysis reported in this paper is restricted to prime-age adult men engaged in regular wage employment.

Trade Protection and Wages

There are at least three channels through which trade protection can affect the industry wage structure: a shock to labor demand, as in traditional trade theories; a change in the product market structure; and a change in industry- or firm-level productivity. Regarding the first of these channels, in short- to medium-run trade models that assume imperfect labor mobility across industries, a reduction in the industry tariff implies a corresponding fall in the relative industry wage. The lack of labor reallocation following episodes of trade liberalization that has been documented for several developing countries (see Goldberg and Pavcnik 2004) lends some support to the notion that adjustment in industries experiencing trade reform occurs via the wage rather than the employment channel. Second, the relaxation of trade barriers induces a pro-competitive effect that reduces the distortionary effects of imperfect competition and erodes industry rents, thereby reducing relative wages. Rodrik (1997) argued that trade increases the own price elasticity of labor demand in absolute terms and thus erodes the bargaining power of labor vis-à-vis capital in the sharing of industry rents. Hasan et al. (2003) found some evidence of a positive relationship between the elasticity of demand for labor and trade protection in India between 1980 and 1997. Third, trade reform also affects industry- or firm-level productivity, though theoretically the direction is ambiguous. There is some evidence from recent studies spanning the 1970s through the 1990s that the trade-productivity link in Indian manufacturing is positive (Das 2002). This paper does not attempt to disentangle these explanations, but merely to examine the empirical evidence for India during a period of rapid trade reform.

One of the earliest attempts to link trade and wage premia was undertaken by Gaston and Trefler (1994) for U.S. manufacturing industries in 1983. They found a statistically significant negative effect of tariffs on relative wages that was robust with respect to the inclusion of industry-specific effects and to the endogenous treatment of tariffs. In contrast, Feliciano (2001), Hasan and Chen (2003), and Pavcnik et al. (2004) found that tariffs, once industry fixed effects were controlled for, had no statistically significant effects in Mexico, the Philippines, and Brazil, respectively. Attanasio et al. (2004), on the other hand, reported a positive correlation between tariff protection and industry wage premia (after controlling for industry fixed effects) that is robust with respect to the inclusion of industry characteristics and instrumenting for trade policy changes. Attanasio et al. (2004), on the other hand, reported a positive correlation between tariff protection and industry wage premia (after controlling for industry fixed effects) that is robust with respect to the inclusion of industry characteristics and instrumenting for trade policy changes. Jean and Nicoletti's (2002) estimation of industry wage premia across 12 OECD countries including the United States using panel data for 1996 revealed a strong positive impact of tariff (and non-tariff) barriers on relative wages in manufacturing industries. Similarly, Arbache et al. (2004) also estimated a positive tariff-wage effect for Brazil between 1987 and 1998.

The handful of recent studies that have focused on India also convey an ambiguous picture of the wage-trade relationship. Dutt (2003) and Goldar (2003) examined the effect of trade protection on average industry product wage (the cost to the employer of hiring workers as opposed to the wage actually received by the worker). To my knowledge, Kumar and Mishra (2004) and Topalova (2004) are the only other studies examining the effect of trade protection on industry wage premia (that is, returns to industry affiliation after controlling for individual characteristics) for India. These studies' findings and points of difference with the current paper are discussed in more detail below.

Labor Market Policy in India

Labor market legislation in India essentially covers workers in the organized sector.[1] The central labor laws are those regulating working conditions (the Factories Act of 1948), employment security (the Industrial Disputes Act of 1947 and the Industrial Employment [Standing Orders] Act of 1946), and trade union activity (the Trade Union Act of 1926).

The Industrial Disputes Act (IDA) is the key piece of legislation governing the relationship between employers and workers. It also provides for a tripartite dispute settlement mechanism, with the state playing a central role. The Trade Union Act regulates the registration and operation of trade unions and allows any seven workers to form a trade union. However, as there is no provision for union recognition and political considerations tend to motivate the position taken by the state, the collective bargaining process is complicated, unions have proliferated, and problems associated with inter-union rivalry have grown (Dutt 2003). As both central and state governments are empowered to legislate on matters relating to trade unions and industrial and labor disputes, firms located in different states might face different regulatory climates (Besley and Burgess 2004). Despite the widespread industrial and trade liberalization during the 1990s, labor reforms proceeded at a very slow pace. In fact, reforms aimed at increasing flexibility with respect to laying off employees, outsourcing, and sub-contracting were introduced only in 2002.

Some economists have argued that labor market regulations combined with the wage-setting system—the system, generally determined by Wage Boards and Pay Commissions in the public sector, that sets the benchmark for private sector wages—and labor redundancy have introduced rigidities in the organized labor market. In response to these rigidities, the argument runs, employers have tended to substitute capital for labor or casual labor for regular labor, or else to subcontract (Ghose 1995). Other observers argue that, due to poor regulatory compliance and spotty enforcement, these laws were not enough of a presence to induce widespread reaction. Moreover, these laws were mostly applicable to the organized sector, which employs only about 8–10% of workers. The empirical evidence on the impact of the IDA regulations on employment growth is mixed: while Fallon and Lucas (1993) reported an adverse impact for the period 1960–82, Dutta Roy (2003) found evidence of rigidities in the adjustment of labor to shocks between 1960 and 1995 but a minimal impact of legislation.

Trade and Industrial Reforms in India

The late 1980s and the 1990s witnessed rapid liberalization of the trade and industrial policy regime (see Nouroz 2001 and Kapila 2001 for a detailed review). The external sector reforms included the liberalization of foreign exchange controls and foreign direct investment. With respect to trade policy, there were substantial reductions in both the level and dispersion of tariff rates, as well as in the numerous general, end-use, specific user, and preferential area exemptions on tariffs. For instance, between 1983–84 and 1999–2000 the peak tariff rates fell from 135% to 39%, while the average tariff rates fell from 98% to 40%. The dispersion between different tariff lines also declined (the standard deviation of unweighted tariffs fell from 30% to 14% during this period). Non-tariff barriers on all goods, other than agricultural and consumer goods, also decreased substantially during this period (Pandey 1998). At the same time, the introduction of five-year export-import (EXIM) policies in 1992–97 and 1997–2002 brought stability and transparency to the system.

During 1980–85, almost 70% of manufacturing industries (with a combined value added share of 45%) were covered by effective rates of protection (ERPs) of between 50% and 150%. Toward the end of the 1990s no manufacturing industry had an ERP in excess of 100%, and about

[1]This comprises all establishments that employ ten or more workers and use electric power as well as those that employ twenty or more workers but do not use electric power.

73% of industries fell within the ERP range of 0–50%. Similarly, there was a drastic fall in non-tariff barriers—while about 92% of manufacturing industries were subject to 100% import restrictions in 1980–85, only 7% were so covered by 1996–2000. The share of manufacturing industries in the zero or minimum import coverage range (0–25%) rose from 8% in the first phase to 72% in the final phase (Das 2003).

Thus, there was rapid and comprehensive trade liberalization during the 1990s, especially with respect to the manufacturing sector. Though India's trade barriers remain higher than those of most developing countries, including post-reform South Asian countries,[2] the liberalization during the 1990s was unprecedented in Indian economic history.

These trade reforms were accompanied by industrial policy reforms with respect to licensing and foreign investment as well as the role of public sector enterprises and large firms The number of industries in which a firm had to obtain a license in order to start production or expand existing capacity was reduced dramatically, as was the number of goods reserved for production by the public sector alone. The reforms also encouraged the inflow of foreign capital by allowing automatic approval in all areas except for a small "negative list," laying down rules for approval in other cases, and simplifying and expediting the procedure (Kapila 2001).

The liberalization process triggered strong GDP growth of about 6.4% per annum during the 1992–2000 period that was accompanied by strong export and import growth. The trade to GDP ratio rose gradually from 15% to 21% between 1990 and 1999 (Ministry of Finance, various years). This paper focuses on the labor market outcomes of regular wage workers, comprising about 25% of the labor force, during this period of rapid structural change. These workers were predominantly employed in manufacturing, public enterprises, and public administration. In line with the strong GDP growth, real weekly wages of these workers grew by about 5% per annum over the same period. At the same time, Table 1 indicates considerable wage dispersion across industries, such as agriculture, trade and hotels, light manufacturing, and other industries that were dominated by private unorganized activity paid the lowest wages, while industries dominated by the public sector, such as mining/quarrying and utilities/services (public administration), paid the highest wages. The latter also experienced a substantial increase in the rate of growth of wages between 1993 and 1999 compared to the previous subperiod (1983 to 1993), possibly due to the increase in public sector wages following the Fifth Pay Commission. Wage dispersion across industries nearly doubled over time, suggesting that industry affiliation could potentially play a role in explaining wage differentials.

Methodology

Following the standard labor economics literature, I estimate wage regression models using augmented Mincerian earnings equations controlling for human capital, industry affiliation, and various other characteristics. The issue of selection bias is addressed using the generalized framework popularized by Lee (1983). Selection is modeled as a polychotomous outcome between three employment categories: non-wage earners (including non-participants in the labor market, the self-employed, and unemployed individuals), regular wage workers, and casual wage workers. As the bias is mediated through observed wages, it is sufficient, and computationally convenient, to divide individuals by employment status, with non-wage earners comprising one group and two different types of wage earners comprising the other two groups.

A two-step model for selection and wage determination is posited. First, a multinomial logit model is estimated where the probability that individual i is in outcome j can be expressed as

[2] As reported in Dean et al. (1994), the average unweighted tariff rate in the early 1990s was 50% in Bangladesh (1993), 65% in Pakistan (1990), 25% in Sri Lanka (1992), and 71% in India (1993).

Table 1. Real Weekly Wages (in 1983 Prices) of Regular Workers by Industry.

	Mean Real Weekly Wage (Rs.)[a]			*Growth per Annum (%)*	
Industry	*1983*	*1993*	*1999*	*1983–93*	*1993–99*
Agriculture and Allied Sectors	53.26	82.16	125.31	6.03	8.76
Mining and Quarrying	177.45	245.02	334.99	4.23	6.12
Light Manufacturing	119.32	145.68	169.22	2.45	2.69
Heavy Manufacturing	161.37	214.12	230.77	3.63	1.30
Utilities	180.16	263.73	423.04	5.15	10.07
Construction	138.04	190.9	223.21	4.25	2.82
Trade and Hotels	89.45	111.58	164.89	2.75	7.96
Transport, Storage, and Communication	150.17	193.6	246.5	3.21	4.55
Services	170.11	249.87	357.92	5.21	7.21
Economy:					
Mean Wage	141.48	206.81	267.03	5.13	4.85
Standard Deviation	112.27	153.18	220.34		

Source: Author's calculations from National Sample Survey employment surveys.

$$(1) \qquad P_{1i} = \frac{1}{1+\sum_{j=2}^{3}\exp(z'_{ji}\gamma_j)}; \text{ and}$$

$$P_{ji} = \frac{\exp(z'_{ji}\gamma_j)}{1+\sum_{j=2}^{3}\exp(z'_{ji}\gamma_j)},$$

$$\gamma_1 = 0;\ j = 2, 3,$$

where the vector z_i comprises a set of exogenous explanatory variables and subscript j denotes the different employment categories. The Theil normalization is applied to the category comprising the non-wage earners, and this category's parameters are thus set to zero to resolve an indeterminacy associated with the MNL model. In addition, the identification of the selection effects is crucial. In the context of the current application, this requires a set of variables that influence employment status but not the wage. These variables (see below) are included as regressors in the selection equation but not the wage equation. The selection correction term, λ_j, is empirically constructed using the MNL estimates from the selection equation and is similar in spirit to the inverse of the "Mills ratio" term.

The second step involves estimation of the wage equation for the j employment sectors. In general, these can be expressed as

$$(2) \qquad w_j = x'_j\beta_j + d'_j\delta_j - \beta^*_j\hat{\lambda}_j + \upsilon_j, \qquad j = 2, 3,$$

where w_j is the vector of natural log of real hourly wages that are observed only for persons engaged in wage employment, the vectors x_j and d_j comprise exogenous explanatory and industry affiliation variables, respectively, and υ_j represents the random error terms such that $E(\upsilon_j|x_j;z_j) = 0$.

This two-step procedure controls for the underlying process by which the set of observations actually observed is generated and provides unbiased estimates of the parameters of the wage equations.[3] The sampling distribution for the estimates is obtained by bootstrapping each of the wage regression models in this paper using 1,000 replications. For brevity, the selection and wage equations are not reported here.[4] The explanatory variables common to both models are worker characteristics such as age, the highest level of education completed, marital status, caste, and religion, as well as controls for settlement

[3]The Lee correction was chosen over other methods of selection correction in polychotomous outcome models because of its simplicity, computational convenience, and transparent interpretation of the selection effect. Using power series approximations for the selection term following the semi-parametric approach advocated by Newey (1999) yields very similar results.

[4]Results are available on request to the author.

type, state of residence, and seasonality effects. As noted earlier, the parameters of the wage equations are identified using a number of exclusion restrictions relating to variables that capture household structure—household size and four dependency variables. The wage regression models include all industry affiliation dummy variables and are thus estimated without a constant term. The estimated models have quite high explanatory power in all three years, accounting for over half of the variation in log wages. The estimated effects are correctly signed and of plausible magnitude, and the majority are significant at the 1% level or better.

The industry fixed effects estimated in the wage equation models are normalized as deviations from an employment-weighted mean differential following Krueger and Summers (1988). The corresponding standard errors are constructed using the procedure suggested by Haisken-DeNew and Schmidt (1997). The resulting wage premia represent the difference between the wage received by an average worker in a given industry and that received by an average worker in the economy. The impact of trade liberalization on the industry wage structure is then examined using a pooled weighted least squares regression model,

$$\text{(3)}\qquad \delta^*_{kt}(\theta_{kt}) = (\theta_{kt})\,T'_{kt}\alpha + (\theta_{kt})\,z'_{kt}\beta + (\theta_{kt})\zeta_{kt}$$
$$k = 1, \ldots, K \text{ industries};$$
$$t = 1983, 1993, 1999,$$

where δ^*_k are the estimated wage premia, T_k represents tariffs, z_k comprise various industry characteristics, and ζ_k is the random error term. Since the dependent variable is constructed from wage equation estimates, a Weighted Least Squares (WLS) estimation procedure that assigns a larger weight to the more precisely estimated industry effects is preferred. The weights ($\theta_k = \sqrt{1}/\text{var}_k$) are the inverse of the variances of the estimated wage premia obtained from the wage equations at time t. The observations are pooled across the three years to enhance efficiency. The analysis is conducted using a unique, specially constructed database for India that combines data from employment surveys with industry-level data. The employment surveys were conducted by the National Sample Survey Organization (NSSO) during January–December 1983, July 1993–June 1994, and July 1999–June 2000 (referred to as 1983, 1993, and 1999).[5] The sample is restricted to men aged between 15 and 65 years engaged in regular wage employment. See the Data Appendix for details.

Industry Wage Premia

The empirical analysis of the links between tariff protection and relative industry wages focuses on a subset of 31 manufacturing industries. This approach is adopted because tariffs do not adequately capture protection in agricultural industries, as these remain subject to numerous quantitative restrictions. Even as late as 1997–98, 84% of value added in agriculture was subject to import licensing requirements, as compared with 30% in manufacturing industries (Cadot et al. 2003). Approximately one-quarter of regular workers are employed in manufacturing industries. The vast majority of regular workers are employed in non-tradable sectors such as public administration and other service industries for which tariff data are not applicable. One manufacturing industry—beverages—was identified as an outlier (see below) and subsequently excluded from the analysis.

The industry wage premia obtained from the wage regression models are generally sizeable and range from a minimum of –24% in the non-motorized and miscellaneous transport equipment sector to a maximum of 26% in the petroleum products sector in 1999. High-wage sectors include heavy manufacturing, such as the petroleum, chemicals, machinery, and transport industries, and low-wage sectors comprise light manufacturing, such as the foodstuffs, tobacco, and textiles industries. An analysis that combines these premia with the industry-level data used in this research shows that high-wage sectors

[5]The employment survey for 1987–88 could not be used because over 76% of the observations on rural wages for persons participating in wage employment are missing.

tend to have a lower share of casual workers, a higher share of skilled workers, and higher capital-labor ratios. These results are consistent with the findings of studies on other countries in which industries that are capital-intensive or skill-intensive (or both) have higher wage premia (Dickens and Katz 1987; Hasan and Chen 2003).

Empirical studies of industry wage structures in the United States and OECD countries have found that these are relatively stable over time (Zanchi 1995), though the evidence for developing countries is mixed (see, for example, Attanasio et al. 2004 and Hasan and Chen 2003). The Indian wage structure in manufacturing industries was relatively stable between 1983 and 1993, though less so between 1993 and 1999 (Spearman's correlation coefficient is 0.877 between 1983 and 1993, 0.809 between 1993 and 1999, and 0.761 between 1983 and 1999). A Kruskal-Wallis equality of populations rank test rejects the null hypothesis that the wage premia estimated in the three years are for the same population (χ^2_2 statistic = 16.90). The overall variability in manufacturing industry wages (summarized by the employment-weighted standard deviation of the industry wage premia adjusted for sampling variance) rose from about 10–11% in 1983 and 1993 to 14% in 1999. The extent to which this change in industry wage dispersion can be attributed to the trade liberalization of the 1990s is assessed below.

Protection and the Industry Wage Structure

The tariff data reveal that the structure of protection across industries also changed somewhat over time: a Kruskal-Wallis test of whether the tariff measurements for the three years could all have been obtained from the same population is decisively rejected by the data (χ^2_2 statistic = 67.19). The Spearman rank correlation coefficient for tariff rates for manufacturing industries is 0.552 between 1983 and 1993, 0.678 between 1993 and 1999, and 0.465 between 1983 and 1999. This is much lower than the correlation coefficient of 0.98 in U.S. tariffs between 1972 and 1988 and of 0.91 in Philippine tariffs over the period 1988 to 1998 (Hasan and Chen 2003), although Attanasio et al. (2004) found that Colombian tariffs have been only loosely correlated over time (correlations range from 0.46 to 0.94 between various year pairs). The relatively weak intertemporal correlation in Indian tariffs is an indication of the extent of trade liberalization experienced during this period.

At the same time, labor reallocation from industries with higher relative tariff reductions was quite low during this period, and industry employment shares changed by less than one percentage point between 1983 and 1999 in all but nine industries. A regression of the change in employment shares on the change in tariff rates suggests that there is no statistically significant relationship between the two. A number of studies have similarly found no evidence of inter-industry labor reallocation in response to trade liberalization in several other countries, including Colombia (Attanasio et al. 2004), Mexico (Feliciano 2001), the Philippines (Hasan and Chen 2003), and Brazil (Pavcnik et al. 2004). This pattern of evidence provides some support for the notion that regular workers were relatively immobile in response to trade shocks (see the section on Indian labor market policy) and that labor market rigidities in developing countries could conceivably restrict labor reallocation following economic reforms.

In order to examine econometrically the impact of the tariff reductions during the 1990s on the industry wage structure, a pooled model is estimated by WLS. The beverages industry was identified as an outlier as this industry had high leverage values and moderately high standardized residual values and excluding these observations materially altered the estimated coefficient on tariffs.[6] A possible reason for the beverages industry being an outlier is that this industry was subject to the highest tariffs in all three years and was also one of the few industries in which tariffs rose between 1983 and 1993. As a

[6]Though some other observations were identified as possible outliers (for example, manufacture of tobacco in 1993), omitting these latter observations from the analysis did not influence the estimated coefficients.

result, the subsequent analysis is conducted for the remaining 31 manufacturing industries; omitting this industry yields a sample of 31 manufacturing industries. In the first specification, wage premia are regressed on the average industry tariff (expressed in fractional points). Dummy variables for the years 1993 and 1999 are included to control for time-specific macroeconomic shocks.

For comparability with other studies, I control for differences in industry characteristics by including 30 industry fixed effects (cotton textiles is the omitted reference category) in the second specification. However, I take the view that, given the small sample size of 31 observations at only three points in time, it would be inadvisable to saturate the model with the complete set of industry fixed effects. Therefore, I also control for inter-industry differences by including industry-specific characteristics in the third specification. These include the worker fatality rate to capture wage differentials reflecting compensation for non-pecuniary aspects of work and the industry shares of skilled and female workers to capture the effects of the average level of human capital and discrimination (Dickens and Katz 1987).[7] The effect of inter-state differences in labor market institutions in federal India is captured by the share of employment concentrated in states where labor market regulations are pro-worker or pro-employer relative to neutral states. Two dummy variables are introduced to capture industrial policy reform with respect to licensing and reservation of production for the public sector. The union threat model argues that firms may pay high wages to avoid the threat of collective action. As a result, higher wages are likely in industries where conditions are favorable to unionization, workers are predisposed to form unions, and firms either have the ability to pay higher wages or—whatever their ability to pay—see a higher percentage of their profits appropriated by labor (Jean and Nicoletti 2002). I therefore also include industry characteristics such as unionization, average establishment size, and capital intensity (Dickens and Katz 1987).

As the predictions of trade theory are predicated in terms of the impact of the *change* in protection on the *change* in relative wages, a differenced model that explains the changes in industry wage premia is also estimated. The explanatory variables are the changes in trade and industrial policy—the change in tariff rates, de-licensing for industries, and de-reservation of production for the public sector. In the final specification, industry dummies for groups of similar industries are included to control for heterogeneous effects.[8]

Table 2 reports the results of the three specifications for the sample of 31 manufacturing industries (which excludes the beverages industry) estimated using WLS.[9] Table 3 reports the results of the differenced models estimated using WLS with the inverse of the sum of the variances in the two years as the weights, as well as those estimated using standard Ordinary Least Squares (OLS) with the White/Huber adjustment for standard errors.

While the first levels specification (without industry controls) explains only about 12% of the variation in the industry wage premia, the third and preferred specification (with indus-

[7]Though the effect of individual productivity or skill (proxied by education) is captured in the wage regression models, the average level of skill in an industry may also influence the average wage. These variables are potentially endogenous; however, the results indicate that the effect of the variable of interest, tariff rates, is robust with respect to the inclusion of these variables.

[8]If industry effects are fixed over time, differencing would eliminate these. However, the first-stage wage equations allow industry effects to differ over time, indicating that it would be appropriate to include industry controls in the second stage.

[9]As the bootstrap is usually designed to construct a confidence interval rather than point estimates, two alternative weights were used as robustness checks: robust variances obtained from wage equations estimated with a standard White/Huber adjustment; and the industry employment share. The results are similar and are not reported here. In particular, the tariff coefficient in the preferred specification (A2) is 0.149 (0.066) using the inverse of the White/Huber adjusted variances and 0.193 (0.054) using the industry employment share compared with 0.207 (0.051) using the inverse of the bootstrapped variances (standard errors in parentheses).

Table 2. Tariff Protection and Industry Wage Premia.
(dependent variable: estimated industry wage premia[a] for manufacturing industries)

Explanatory Variables	*(A1)*	*(A1')*	*(A2)*	*Mean*
Trade Policy:				
Average Tariff[a]	0.141*	0.105*	0.207***	0.713
	(0.080)	(0.055)	(0.051)	(0.325)
Compensating Differentials:				
Frequency Rate of Fatal Injuries			0.341	0.058
			(0.243)	(0.039)
Labor Quality:				
Share of Female Workers			–0.236***	0.089
			(0.084)	(0.121)
Share of Skilled Workers			0.436***	0.418
			(0.049)	(0.187)
Labor and Industrial Policy:				
Pro-Worker States			0.083	0.434
			(0.087)	(0.164)
Pro-Employer States			0.138	0.364
			(0.101)	(0.147)
Industrial License Dummy			–0.005	0.484
			(0.029)	(0.502)
Public Sector Dummy			0.027	0.129
			(0.021)	(0.337)
Organized Sector Variables:				
Union Density			0.060	0.186
			(0.040)	(0.157)
Average Establishment Size			0.0003***	91.69
			(0.0001)	(112.230)
Capital-Output Ratio			0.074***	0.361
			(0.028)	(0.242)
Constant	–0.128	–0.114**	–0.477***	
	(0.084)	(0.055)	(0.096)	
Time Dummies	Yes	Yes	Yes	
Industry Dummies	No	Yes	No	
Number of Observations	93	93	93	93
R-Squared [b]	0.115	0.872	0.693	

Notes: The beverages industry has been excluded from the dataset. All regression models are estimated using Weighted Least Squares with the inverse of the variance in the estimated wage premia as weights.

[a]Wage premia and tariffs are expressed as fractional points. The mean wage premium for the sample is –0.018.

[b]The R^2 for the WLS models is computed as the squared correlation coefficient between the actual and predicted values of the dependent value.

*Statistically significant at the .10 level; **at the .05 level; ***at the .01 level.

try-specific variables) explains almost 70% of this variation.[10] The goodness of fit measures reported for the differenced models are, not surprisingly, considerably lower.

[10]Conventional Chow tests of pooling restrictions are not valid, as the models are estimated using weighted least squares. Instead, a Wald test using the weighted variance-covariance matrix from the regression model was used to test the joint significance of the interaction terms. The null hypothesis that the coefficients on the interaction terms are not statistically significantly different from zero could not be rejected in any of the specifications (χ^2_2 = 1.85 for specification A2).

Table 3. Tariff Protection and Industry Wage Premia: Differenced Model. (dependent variable: change in the estimated inter-industry wage premia for manufacturing industries)

	WLS Models[a]			*OLS Models*[b]			
Explanatory Variable	*(D1)*	*(D2)*	*(D3)*	*(A1)*	*(A2)*	*(A3)*	*Mean*[c]
Change in Tariff	0.193***	0.208***	0.269***	0.088	0.105	0.182**	–0.352
	(0.056)	(0.059)	(0.061)	(0.073)	(0.072)	(0.087)	
Delicensing[d]	0.006	–0.020	–0.029	0.021	0.003	–0.014	0.436
	(0.017)	(0.024)	(0.024)	(0.020)	(0.023)	(0.022)	
Public Sector De-Reservation[d]		0.0269	0.039*		0.0245	0.031	0.097
		(0.0220)	(0.023)		(0.0207)	(0.020)	
Constant	0.004	0.005	0.061*	–0.039	–0.037	0.037	
	(0.025)	(0.025)	(0.036)	(0.035)	(0.034)	(0.078)	
Time Dummy	Yes	Yes	Yes	Yes	Yes	Yes	
Industry Group Dummies	No	No	Yes	No	No	Yes	
Observations	62	62	62	62	62	62	
R^2	0.091	0.140	0.382	0.067	0.090	0.225	

Note: The beverages industry has been excluded from the dataset.
[a]These models are estimated using Weighted Least Squares with the inverse of the sum of the variances in the two years as the weights.
[b]These models are estimated using Ordinary Least Squares with White/Huber corrected standard errors.
[c]The mean of the dependent variable is –0.059.
[d]These variables are coded zero if there is no change and one if there is a change between two years. All change implies a removal of licensing and reservation.
*Statistically significant at the .10 level; **at the .05 level; ***at the .01 level.

Tariff-Wage Elasticity

The tariff-wage effect is consistently positive and significant at the 10% level or better across all three specifications, indicating that industries with relatively higher protection would tend to have relatively higher wages. The estimated coefficient is more precisely determined as additional industry covariates are included and becomes significant at the 1% level or better. As mentioned earlier, I prefer to capture industry effects using industry-specific characteristics rather than using 30 industry fixed effects. However, I find that the tariff-wage elasticity is positive, statistically significant, and of broadly similar magnitude regardless of how industry effects are controlled for.[11] The third and preferred specification with industry-specific variables (A2) reports a tariff-wage elasticity of 0.21. This implies, for instance, that moving a worker from an industry with the average level of protection in 1983 (a 103% tariff rate) to an industry with the average level of protection in 1999 (40%) would lead to a substantial fall of about 13% (= 0.207 ∗ (1.03 – 0.40) ∗ 100) in the average worker's wage. The differenced models also report a positive and statistically significant relationship of a similar magnitude between the change in tariffs and the change in wages—that is, industries with larger tariff reductions tended to have larger reductions in wages.

As discussed earlier, the existing empirical evidence is mixed. Gaston and Trefler found a negative correlation between tariff protection and industry wage premia in the United States, while other studies for Brazil (Pavcnik et al. 2004), Mexico (Feliciano 2001), and the Philippines (Hasan and Chen 2003) have estimated a non–statistically significant tariff-wage elasticity. In contrast, Jean and Nicoletti (2002) estimated a positive tariff-wage effect

[11]Following the other studies cited in this paper, I also estimate the third specification (not reported) with both industry-specific variables and industry fixed effects. The estimated tariff coefficient of 0.118 (0.068) is very similar to that obtained in the first and second specifications.

after controlling for industry fixed effects for a panel of OECD countries (including the United States). Similarly, Arbache et al. (2004) estimated a tariff-wage elasticity of 0.16 (evaluated at the median tariff) for Brazil between 1987 and 1998, and Attanasio et al. (2004) obtained a positive tariff-wage elasticity of 0.14 for Colombia. Both of the latter estimates are similar to the estimated tariff effect reported here in Tables 2 and 3.

In common with the evidence for other countries, the evidence for India has been ambiguous. Two studies, Dutt (2003) and Goldar (2003), examined the effect of trade protection on *average* industry product wage for the organized manufacturing sector using firm or industry data from the Annual Survey of Industries (ASI). Two other studies, Kumar and Mishra (2004) and Topalova (2004), are closer in spirit to this paper in that they examined the impact of trade protection on *relative* industry wage premia using household survey data from the NSS. Interestingly, the two studies within each of these pairs obtained contradictory results, despite basic similarities in the data source and methodology.

For instance, both Dutt and Goldar estimated a fixed-effects panel model using ASI data, but whereas the former detected no statistically significant relationship between nominal tariff protection and real average product wages, the latter found a statistically significant positive impact of effective tariff protection and non-tariff barriers on the real average product wage. The key difference between these two studies was the definition of tariff protection and the time period (Dutt's study covered three years during the 1990s, while Goldar's covered 18 years from 1980–81 through 1997–98).

At the same time, both studies differ from this paper with respect to wage variable definition and source, methodology, and sample coverage. Dutt and Goldar focused on the average real annual product wage, whereas this paper focuses on the relative real hourly consumer wage. The former is constructed by dividing the reported industry annual wage bill by the number of employees (male and female) and is applicable to the organized manufacturing sector only. The latter, in contrast, is used to estimate industry wage premia—that is, wage differentials arising solely out of the worker's industry affiliation, where the wage variable is the weekly wage rate reported by all male regular wage workers in the NSS household surveys divided by the number of hours worked during the week. None of the NSS surveys except the last round report whether a worker is employed in the organized or unorganized sector. However, the main reason Dutt's findings differ from those of both the present study and Goldar (2003) appears to be the time period covered: Dutt examined only the 1990s (post–trade liberalization), as opposed to both the 1980s and 1990s (pre- and post–trade liberalization).[12]

Similarly, the findings of the only two other studies that investigate the effect of trade protection on relative industry wages (after controlling for individual characteristics) for India are also contradictory. Kumar and Mishra (2004) obtained a statistically significant negative effect of tariffs on industry wage premia for urban workers across all specifications (levels and differenced) that was robust with respect to instrumenting for tariffs. In contrast, Topalova (2004), who estimated a levels specification with lagged tariffs, industry controls, and time controls as explanatory variables for wage premia constructed for urban and rural workers separately, found a statistically significant positive tariff-wage elasticity of about 0.14 to 0.16 in the former (the point estimate for the elasticity for the rural sample was similar in magnitude but not statistically significant).[13]

While both Kumar and Mishra (2004) and Topalova (2004) employed a methodology similar to that in this paper, there are im-

[12]Estimating specification (A2) using OLS and with the natural log of the mean industry wage as the dependent variable yields a statistically significant positive tariff-wage elasticity of 0.182 (0.101) that does not materially differ from the estimate obtained with the wage premia as the dependent variable. However, the tariff-wage elasticity is no longer significant at any conventional statistical level when this OLS model is estimated only for the two years in the 1990s.

[13]Topalova also found a comparable positive tariff-wage effect using average industry wages obtained from ASI data for the organized manufacturing sector.

portant differences with respect to sample coverage and the earnings measure used. Kumar and Mishra obtained their industry wage premia from wage equations estimated over a sample of all urban workers, male and female, engaged in wage employment and self-employment in manufacturing industries (other than food processing industries) using NSS survey data for 1983, 1987–88, 1993–94, and 1999–2000. Topalova used separate urban and rural samples of male and female wage workers in agricultural, mining, and manufacturing industries for 1983, 1993–94, and 1999–2000. It is debatable whether the same wage determination process applies across these different categories of workers, especially given the low female labor force participation rate in India (about 30% of adult women aged 15–65, compared to about 85% of adult men). In addition, only 2–3% of workers classified as self-employed report non-zero wages in the NSS survey data, suggesting that including these workers with the sample of wage workers is problematic.

The other key difference is the earnings measure used: Kumar and Mishra (2004) used real weekly wages with no controls for hours worked in the wage regression models. This might be important given that some of their wage variation across industries may be attributable to hours variation. The present study, and the other studies cited earlier, all use real hourly wages.[14] There are also some other differences in variable definition and in the specification of the first-stage regression models.[15]

[14]Though using the constructed real hourly wage variable (see the data appendix) introduces measurement error in the dependent variable, this would be captured in the error term in the wage regression models. On the other hand, using the reported real weekly wage as the dependent variable without controlling for the hours worked or including this as a control might lead to biased parameter estimates due to omitted variable bias or because of correlation of measurement error with the error term.

[15]For instance, Kumar and Mishra included controls for occupation in the wage equation. There is no consensus in the literature regarding the inclusion of such controls, as they could be legitimately interpreted as endogenously determined with wages. In addition, none of the above authors address the issue of selection into wage employment.

The positive relationship between tariff protection and relative wages reported in this paper is consistent with the notion that the lack of competition from imports due to high tariff barriers raises the demand for labor, which in turn raises relative industry wages as labor reallocation across industries is constrained and does not eliminate the differential. The existence of labor market rigidities seems relevant in the Indian context, where employment security is of paramount importance for workers (Ghose 1995) (see also the section on Indian labor market policy). Data from the 1999 NSS survey reveal that there is practically no mobility of workers across industries—only about 1% of all workers and 1.2% of male workers had changed industries (at the two-digit NIC level) in the two years preceding the survey. Presumably workers with regular jobs build up firm- and industry-specific human capital or acquire seniority status that makes their response to labor demand shocks less elastic. This positive tariff effect is also consistent with the notion that trade reduces distortions in an imperfectly competitive market, eroding rents and leading to a fall in relative wages.

Other Industry Characteristics

The estimated effects for the industry and labor policy variables are poorly determined, possibly due to the fact that these are relatively crude measures that fail to adequately capture the differences in the constraints imposed by the policies across industries.[16] In addition, there is no inter-industry variation in the license dummy in 1983, as all manufacturing industries had to obtain licenses for entry and expansion at that time. However, the differenced model (see Table 3) finds that the de-reservation of production by the public sector raised the relative industry wage by about 4%, indicating improvements in efficiency and productivity following deregulation (Jean and Nicoletti 2002).

[16]Data on alternative measures such as the proportion of industry production subject to licenses or public sector production are not available at a sufficient level of disaggregation.

Union density exerted no independent effect on relative industry wages. This result could be taken as indicating that Indian workers value employment security over wages (Ghose 1995). However, the poorly determined nature of the estimated effect is also possibly a consequence of using average industry-level data (given the absence of individual-level data),[17] or under-reporting the true extent of unionization within particular industries, or both.[18] The estimated coefficients on the other control variables are generally plausible. The effects of female and skilled workers on relative industry wages are both well determined and are negative and positive in sign, respectively, as anticipated (see Dickens and Katz 1987). There is some evidence in support of the notion that industries with large average establishment sizes pay higher wages (Brown and Medoff 1989). The time effects are not statistically significant at conventional levels in the majority of the specifications, indicating that time-variant and industry-invariant factors were not statistically significant determinants of wage premia.

Robustness Checks

I now test the sensitivity of the tariff-wage effect to the use of alternative tariff measures and to the inclusion of trade variables.[19] Some researchers prefer to use a measure of the published tariff that does not take into account any exemptions (Pandey 1998), while others use a measure that takes into consideration exemptions that apply to entire tariff lines (Nouroz 2001). The former is preferable if the exemptions on tariffs are included in profits so that domestic prices reflect the full tariff; the latter if these exemptions are incorporated in the domestic prices. The latter measure for average tariff is used in the regression models above. The regression models reported in Table 2 are re-estimated using the alternative no-exemption tariff measure. The tariff-wage effect is still statistically significant and positive, though the elasticity (0.1) is lower than that estimated above.

I also estimate the regression models reported in Table 2 with alternative trade flow variables as additional regressors. These capture the various channels besides tariffs through which trade policy affects wages. The empirical approach adopted here is to separately include alternative measures for (i) contemporaneous real import and export flows in 1983 prices, (ii) import and export shares, and (iii) import penetration and export intensity. As these trade variables are potentially endogenous (since they depend on factor costs), they are treated merely as conditioning variables to test the robustness of the tariff coefficient. The coefficient on tariffs is robust with respect to the inclusion of these trade variables, and the tariff-wage elasticity (0.2) for the preferred specification (A2) is very similar to that reported in Table 2. The estimated coefficients corresponding to the other control variables remain highly stable when these different trade measure proxies are tested in turn.

Endogeneity of Tariff Protection

An important consideration in such research is the potential for a correlation in the unobservables that determine tariffs and relative wages due to political economy considerations, unobservable worker characteristics, or both. To the extent that these factors are time-invariant, they can be addressed, as above, using industry fixed effects, industry-specific characteristics, or differencing (Attanasio et al. 2004). However, if these factors vary over time, a two-stage least squares procedure (or the equivalent) that allows for the simultaneous determination of tariffs and wage premia should be used. The difficulty, commonly acknowledged in the literature, lies in finding appropriate instruments.

Gaston and Trefler (1994) instrumented for tariffs in a cross-section setting using

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[17]Teal (1996) estimated a statistically significant union-wage elasticity ranging from 0.21 to 0.25 using individual-level data for Ghana in the early 1990s. When firm-level data were used, this value fell to 0.07 and was no longer a statistically significant determinant of wages.

[18]Not all unions are legally required to submit annual statutory returns, and many state governments do not publish data on registered trade unions.

[19]Results are available on request to the author.

one set of instruments motivated by reference to the endogenous trade protection literature and a second set capturing the average industry composition of workers. In the absence of a theoretical model of the dynamics of the political economy of trade protection, some of the other studies cited earlier have used instruments based on the history of tariff protection in the relevant country. Thus, Attanasio et al. (2004) and Pavcnik et al. (2004) used the interaction of exchange rate and world coffee prices with the pre-sample tariff rate to instrument for tariff changes. Similarly, Kumar and Mishra (2004) instrumented for tariffs using interactions of the foreign exchange reserves with the tariff rate in 1980 and with the share of unskilled workers in 1983.

In this paper I attempt to instrument for tariffs by estimating a feasible efficient two-step Generalized Method of Moments (GMM) estimator in which the optimal weighting matrix is formed by the IV residuals. I estimate six regression models using a range of instruments suggested both by the endogenous protection literature and by the evolution of protection in India.[20] The first two sets of instruments are based on Gaston and Trefler's models: variables suggested by endogenous trade policy (such as exports, imports, scale, capital stock, share of casual workers, and share of production in the organized manufacturing sector); and average worker characteristics (age, education, caste, religion, marital status) and the share of casual workers. There is some evidence that trade liberalization in India progressed to some extent according to the end-use classification of goods (Nouroz 2001). As a result, I use the share of employment in industries (at the three-digit NIC level) engaged primarily in the production of consumer goods within the broader industry category used in this research as an instrument for tariff rates along with the interaction of 1983 tariffs with either the real exchange rate or foreign exchange reserves. The final two sets include the share of intermediate sales in total domestic output (Cadot et al. 2003) derived from input-output tables along with either of the two interaction terms.

An obvious requirement with this approach is that the identifying instruments be highly correlated with the endogenous variable and orthogonal to the error process governing the structural equation. In particular, an F-test of the joint statistical significance of the excluded (or identifying) instruments from the first-stage reduced form equation provides the basis for a test of the first condition (see Bound et al. 1995). The second condition can be investigated through a test of over-identifying restrictions, such as Hansen's J-statistic, which tests the null hypothesis of correct model specification and orthogonality of the instruments with the structural equation's error process.

In only two cases do the instruments appear to be valid in statistical terms. For the base model (A1') and using the third and fourth sets of instruments, the F-tests exceed the "rule-of-thumb" value of 10 (Staiger and Stock 1997) and Hansen's J-statistic is not statistically significant. However, I view the base model (A1') as potentially over-parameterized, given the relatively small sample size and the inclusion of 30 industry dummies. Given the low degrees of freedom, there is potential for the generation of unreliable estimates, and the corresponding statistical tests have weak power. Thus, the model is clearly sensitive to the instruments specified and industry controls used. The Hansen J-statistic for over-identifying restrictions and the "rule-of-thumb" F-values for the preferred specification (A2) suggest that the GMM estimates are likely to be unreliable. It could be reasonably argued that the OLS estimates are likely to provide a more sensible benchmark for the tariff-wage elasticity than the GMM procedures.

There are also other reasons for believing that the potential endogeneity of tariffs in such models is a less serious consideration in the Indian context. Gang and Pandey (1996) found in their empirical analysis for 1979–80, 1984–85, and 1991–92 that neither political factors (lobbying) nor economic factors (the infant-industry argument) were statistically significant determinants of manufacturing tariff levels. They attributed this finding to

[20]Results are available on request to the author.

a possible hysteresis in trade policy since the 1960s, with political factors continuing to ensure a relatively unchanged inter-industry structure of protection despite economic arguments in favor of liberalization.

Similarly, Topalova (2006) found in her study using annual tariff data between 1987 and 2001 at the six-digit HS level that changes in tariffs were uniform at least until 1997. The notion that tariff protection is determined by relative efficiency considerations (as captured by productivity) is not supported by the data at least until 1997. She also found, using ASI data from the organized manufacturing sector, that changes in tariffs between 1987 and 1997 were uncorrelated with variables that capture political protection, such as employment and output size, average wage, average establishment size, and the industry's share of skilled labor. The same result held using NSS data on the average industry per capita expenditure, wage, and indicators of poverty incidence, depth, and severity.

Topalova's findings and Gang and Pandey's analysis are consistent with my largely unsuccessful attempt to obtain appropriate instruments for tariff protection—it appears that none of the standard economic or political factors play a statistically significant role in determining trade policy, while data limitations prevent us from examining in greater detail the specific evolution of tariff protection in India. Given these limitations, I believe that it adds little to our analysis to present potentially dubious IV estimates based on invalid instruments (see also Bound et al. 1995).

Conclusion

This study provides one of the first attempts to analyze econometrically the link between trade protection and industry wage premia (relative industry wages after controlling for worker characteristics) in India. I have found that the impact of trade protection on the industry wage premia among regular workers was statistically significant, with more protected industries tending to pay higher wages relative to the average wage across all industries. However, the magnitude of the impact was small. This positive, albeit modest, tariff-wage effect is robust with respect to the inclusion of industry controls in the form of either industry fixed effects or industry-specific characteristics. This positive effect could derive from the erosion of rents that are received (and are reflected in the wages earned) by unionized workers in imperfectly competitive markets following trade liberalization or returns to industry-specific skills when there is limited labor reallocation. A caveat is that the unavailability of ideal valid instruments rendered the treatment of tariff endogeneity less persuasive than I would have liked. Nonetheless, this research provides some important insights into how India's reforms during the 1990s affected the distribution of wages in manufacturing industries. It also complements the modest though expanding literature on the empirical relationship between trade protection and industry wages.

Another caveat is that of the three external sector reforms during the 1990s in India—the reduction of tariff barriers, non-tariff barriers (NTBs), and barriers to foreign direct investment—this paper has examined only tariff reductions, due to the lack of comparable and accessible data on the latter types of reforms. However, I find that, for the years for which comparable data exist, both the level of and change in NTBs and tariffs tend to be positively correlated. Furthermore, the exclusion of foreign direct investment from the analysis seems unlikely to have materially weakened its conclusions, since the actual inflows were relatively small and accounted for only about 0.4% of GDP between 1990 and 1998 (Rajan and Sen 2001); by comparison, trade represented 17% of GDP during this period (Ministry of Finance, various years).

The findings of this study suggest at least two potentially fruitful areas for future research. First, ideally the short-run costs of adjustments following trade reform should be taken into account in order to examine wider issues of worker welfare. Second, although there are reasons to believe this is less of an issue for India, a study that uses better disaggregated data at the industry level could more adequately address the potential problem of endogeneity of tariffs.

The results of this study suggest that India's trade reforms during the 1990s had implications for wage inequality. The wage regression models indicate that the skill premium for graduates rose during this period. In addition, the data suggest that industries with high initial levels of tariff protection (that is, in 1983) had low wages relative to the economy average and a high share of unskilled workers. These industries were also the ones that experienced the greatest reductions in tariff protection. As a result, the positive tariff-wage effect implies that the trade reforms are likely to have exacerbated overall wage inequality as the relative wages of workers (predominantly unskilled) in these manufacturing industries fell. Unskilled workers were hit by the rising skill premium as well as an additional decline in their relative wages as the industries in which they were predominantly employed experienced a decline in wage premia relative to skill-intensive industries. This is reflected in the rise in the Gini coefficient of real hourly wages from 0.396 in 1983 to 0.439 in 1999 (Dutta 2005). At least some of this rise in inequality is likely to be attributable to the trade liberalization during the 1990s. A decomposition of the wage regression functions using the Fields (2002) methodology reveals that the industry affiliation of workers explains about one-fourth of the wage inequality in each of the three years (Dutta 2005). To the extent that the level of, and change in, trade protection affect the industry wage premia, they are also likely to affect wage inequality.

Low mobility between industries in India, due to the lack of transferable skills or other barriers, prevents workers from moving out of industries with declining relative wages in response to trade reform. This suggests the need to increase labor market flexibility through labor market and other institutional reforms. These reforms would need to be supplemented by adequate provisions for social protection. Safety-net programs for workers affected by trade are necessary to minimize the short-run adjustment costs faced by trade-displaced workers. There is a need for a coherent strategy for social protection in this context such as the rationalization of severance pay schemes, a movement toward insurance mechanisms covering both the organized and unorganized sectors, and skill development programs for workers.

Data Appendix

Variable	*Description*
First-Stage Wage Regression Models (source: National Sample Employment-Unemployment Surveys):	
Employment Status	Individuals were divided into three mutually exclusive categories using current weekly status: (i) non-wage earners, that is, non-participants in the labor market, self-employed individuals, and unemployed individuals; (ii) regular wage workers; and (iii) casual wage workers.
Hours Worked	Constructed from the intensity of work variable for each day of the week preceding the survey. This variable assumes one of three values—no work, part-time (if worked between one and four hours during the day), or full-time (if worked more than four hours during the day). The daily number of hours worked is coded zero if no work, four hours if part-time work, and eight hours if full-time work. This is aggregated for the week, subject to a maximum of 48 hours per week.
Real Hourly Wage	The distribution of nominal weekly wages was trimmed by 0.1% at the top and bottom tails. Nominal wages were then deflated to 1983 prices using official state-level monthly consumer price indices for agricultural laborers (CPIAL) for rural wages and industrial workers (CPIIW) for urban wages. Real hourly wages are constructed by dividing the real weekly wage by the number of hours worked in the week.

Continued

Data Appendix continued

Variable	*Description*
Industry Affiliation	Constructed from the current weekly industry variable reported at the three-digit National Industrial Classification for 1983 and 1999 and five-digit NIC for 1999 and further aggregated into 54 industries to ensure at least ten observations per industry. Concordances between the three different industrial classification systems were constructed.
Highest Level of Education Completed	Four mutually exclusive binary variables for primary, middle, secondary, and graduate schooling—all coded 1 if the individual has completed the given education level, 0 otherwise. The fifth reference category is individuals with zero or less than two years of education.
Age Splines	Five age splines with nodes (or knots) at ten-year intervals.
Married	Coded 1 if currently married and 0 if never married, widowed, divorced, or separated.
Social Exclusion	Three mutually exclusive binary variables: (1) coded 1 if household belongs to scheduled caste or tribe, 0 otherwise; (2) coded 1 if the household is Muslim, 0 otherwise; (3) all others.
Seasonality	Dummy variables for the quarter of the interview; also interacted with the rural dummy.
State of Residence	Dummy variable for state of residence .
Rural	Dummy variable for rural/urban settlement type
Second-Stage Wage-Trade Regression Models:	
Tariff Rate	Simple average tariff rates (including the auxiliary tariff in 1983–84 and the customs surcharge in 1999–2000). The majority are *ad valorem* rates—the *ad valorem* part of composite rates is included while specific rates are not. Those exemptions that are quantifiable and are applicable to all goods within a tariff line are taken into account. Concordances between the different tariff classification systems (the Brussels Tariff Nomenclature in 1983 and the Harmonised System (HS) in the 1990s) and the industrial codes were constructed. Import-weighted tariffs could not be constructed due to unavailability of import data at a sufficient level of disaggregation. Source: Indian Tariff Schedules.
Import and Export Shares	Total value of imports (exports) in the industry as a proportion of total imports (exports) across all industries. Source: Input-Output Tables (Central Statistical Organisation).
Industrial License Dummy	Coded one if industry has to obtain a license for any aspect of economic activity, zero otherwise. Source: Kapila (2000).
Public Sector Dummy	Coded one if industry reserved for the public sector, zero otherwise. Source: Kapila (2000).
Labor Policy Dummies	The share of total industry employment in pro-worker, pro-employer, and neutral states (classified according to the nature of amendments to the Industrial Disputes Act 1947). Source: NSS surveys and the Besley and Burgess dataset (2004).
Share of Skilled (/Female) Workers	Proportion of skilled (/female) workers in regular wage employment in the industry. Source: NSS surveys.
Organized Sector Variables Only (source: Annual Survey of Industries and Labour Bureau publications):	
Capital-Output Ratio	Capital-Output Ratio prices by the wholesale price index (WPI) for machinery, transport equipment, and construction materials divided by output.
Establishment Size	Total number of employees divided by number of factories.
Fatality Rate	Number of deaths as a result of an accident per hundred thousand man-days worked.
Union Density	Membership of trade unions divided by number of employees.

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