Trade Liberalisation and the Industry Wage Structure in India

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Abstract

This paper represents one of the first attempts at analysing econometrically the link between trade protection and inter-industry wage premia in India. This analysis combines detailed tariff data with micro survey data for three years that span the period of rapid trade liberalisation in the 1990s. Augmented Mincerian earnings equations are estimated using a set of human capital measures and a variety of worker, industry and state characteristics after correcting for potential selection bias. Inter-industry wage premia are obtain as deviations from an employment-weighted mean differential. This paper finds that the impact of trade liberalisation on the inter-industry wage premia for regular workers is substantial and that industries that undergo tariff reductions have lower wages relative to other industries. This positive tariff-wage effect is evident whether or not industry fixed effects are included and is consistent with the short-run specific factors and the medium-run Ricardo-Viner models of trade.

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1. Introduction

During the 1980s India was described as "one of the most complicated and protectionist regimes in the world" (International Monetary Fund, 1998). Following the macroeconomic crisis in 1991 there was rapid trade liberalisation and domestic deregulation. Despite considerable debate concerning the possible impact on the Indian economy of these reforms little systematic empirical work has been undertaken on the effects of such liberalisation on wages in India. This paper attempts to fill this gap through an econometric examination of the link between trade liberalisation and inter-industry wage premia.

Most empirical work examining the link between trade and wages has focussed on the returns to skill or education and the impact of trade liberalisation on wage inequality between skilled and unskilled workers within the context of the Hecksher-Ohlin framework (Katz and Murphy, 1992; Robbins, 1996). Goldberg and Pavcnik (2003) argue that in developing countries where labour market rigidities prevent labour reallocation across sectors in response to trade liberalisation and where markets have been recently liberalised, short- and medium-run trade models and trade models with imperfect competition that allow for sector-specific returns to factors are more appropriate. As a result the industry affiliation of the worker becomes an important determinant of the wage, either in the form of returns to industry-specific skills that cannot be transferred in the short- to medium-run or as industry rents arising out of imperfect competition. This paper also draws on the literature on inter-industry wage premia as there is some evidence of rigidities in the Indian labour market implying that an industry based approach is appropriate (Tendulkar, 1998).

The next section describes the analytical framework underlying this analysis. Section 3 outlines India's trade policy reforms undertaken during the 1990s. The next two sections detail the methodology and the data used in this paper. This study exploits three national employment surveys 1983, 1993-94 and 1999-2000 that recently have been made available to researchers. The first survey can be interpreted as providing insights into the structure of Indian labour markets prior to liberalisation while the latter two provide the basis for delineating a portrait of these structures after the radical trade liberalisation process. The two-stage methodology of Krueger and Summers (1988) is employed to filter out the effects due to observable worker characteristics from the inter-industry wage premia. The role of trade policy in determining these estimated wage premia is then assessed. Section 6 presents the empirical results for the wage regression models and the regressions of inter-industry wage premia on trade policy and other determinants. The empirical analysis reported in this paper is restricted to prime-aged adult males engaged in regular wage or salaried employment. Section 7 concludes.

2. Analytical framework

The role of trade liberalisation in determining inter-industry wage premia can be analysed along two lines: a shock to the demand for labour and a change in the product market structure.

In the first scenario trade liberalisation is viewed as a shock to industry demand. The product price changes accompanying trade liberalisation result in changes to the composition of output and hence, in the bundle of factors used in production. This will result in changes in wages, if labour supply is fixed, and in employment, if labour supply is flexible. The link from trade liberalisation to relative wages has usually been explored within the Heckscher-Ohlin-Samuelson (HOS) framework. This model predicts that tariff reductions leading to a fall in the relative price of a commodity will reduce the real returns to the factor used intensively in the production of that commodity and increase the real return to the other factor unambiguously (Markusen et al., 1995). Most empirical research has focussed on analysing the impact of trade reform on this single economy-wide return to labour (see for example Robbins (1996), Katz and Murphy (1992)).

The HOS model is essentially a long-run phenomenon that assumes perfectly competitive and integrated markets and complete factor mobility. In the short-run these conditions are unlikely to hold. In addition, as Lang et al.(1987, pp. 4) point out, “labour does not compete in a single aggregate labour market.” Empirical evidence suggests that wages received by apparently similar workers differ across different industries and that this difference arises out of the worker’s industry affiliation even after controlling for ability and other worker characteristics (Krueger and Summers, 1988). Though inter-industry wage premia arising due to worker heterogeneity, compensating differentials and temporary industry demand or supply shocks are consistent with the HOS framework there is considerable evidence of non-competitive explanations for the existence of inter-industry wage premia that are not consistent with this model. Some of these efficiency wage explanations are incorporated into the empirical model discussed in Section 6.4.2. In the HOS model since labour is assumed to be fully mobile in a perfectly competitive world relative industry wages would not be affected by trade liberalisation.

The medium run Ricardo-Viner (RV) model of trade allows imperfect factor mobility with one factor mobile across sectors while the other is taken to be sector-specific. Following a fall in the price of a good due to trade liberalisation the factor specific to the sector that experienced the price reduction loses while the other specific factor gains in real terms. The impact on the real returns to the mobile factor is ambiguous - the real returns fall in the expanding sector and rise in the contracting sector so that the net effect depends on consumers' preferences for the two goods (Markusen et al., 1995). If there are barriers to labour mobility across sectors then this model predicts a positive relationship between protection and industry wage premia – the fall in trade barriers in a sector will adversely affect the relative wage earned by the workers in that sector.

In the second case, trade liberalisation influences the product market structure. The relaxation of trade barriers induces a pro-competitive effect. In the presence of scale economies the number of firms in an industry are limited and there is imperfect competition. By creating a larger market that is capable of supporting a greater number of firms trade liberalisation increases market competition and reduces the distortionary effects of imperfect competition (Markusen et al., 1995). In the presence of imperfect competition and unionisation wages are functions of the firm's product market rents and the worker's reservation wage and the share of the rents appropriated by workers depends on their bargaining power. Those industries with relatively low labour share and greater market power (i.e. those that are more concentrated, face barriers to entry, make higher than average profits) tend to pay higher wages (Jean and Nicoletti, 2002). These industry rents are eroded with trade liberalisation. Rodrik (1997) argues that trade increases the own price elasticity of demand for labour[2] that erodes the bargaining power of labour vis-à-vis capital in the sharing of industry rents. Hasan et al (2003) find that the elasticity of demand for labour in India is positively related to trade protection in the period 1980 to 1997.

[2] By increasing the availability of raw, intermediate and finished goods trade raises the elasticity of substitution between labour and other factors of production as well as the elasticity of demand for the finished good. The first impacts the demand for labour through the substitution effect while the second through the Hicks-Marshallian laws of factor demand (Hasan et al., 2003).

Trade reform also impacts firm productivity, though the direction theoretically (and empirically in the Indian case (Epifani, 2003)) is ambiguous. If there is a positive impact and these enhancements in productivity are passed on to wages within industries then the link between wages and trade liberalisation would also be positive (Goldberg and Pavcnik, 2003).

Thus, despite some ambiguity the empirical relationship between trade liberalisation and relative industry wages seems to be mostly positive. This paper draws on the inter-industry wage premia literature and assesses whether the impact of India's tariff reductions during the 1990s conform to the predictions of the trade models described here. The earliest attempt to link trade and wage premia was made by Gaston and Trefler (1994) for US manufacturing industries in 1983. They find a negative correlation between tariff protection and industry wage premia that is robust to the inclusion of industry fixed effects and to the treatment of tariff as endogenous. Possible explanations for this negative correlation, supported by the data, include the possibility that unions take advantage of protection by offering wage concessions in exchange for employment guarantees. Other explanations could be that long-term protection prevents the efficient reallocation of resources from import-competing sectors (Gaston and Trefler, 1994). Goldberg and Pavcnik (2003) undertake a similar analysis for Columbia and find that after controlling for industry fixed effects trade protection tends to be positively associated with wages.[3] In the Indian context, to the author's knowledge, this is the first attempt to estimate inter-industry wage premia and quantify the role of trade in determining these premia.

A few studies have estimated wage regression models using employment survey data in India. Kingdon and Unni (2001) have estimated wage regression models for male and female workers for the urban sample for two states in 1987-88 and focus on education effects. Duraisamy (2002) estimates the rate of return to education for all male and female workers using wage regression models for 1983 and 1993-94. A comparably specified version of Duraiamy's wage regression model estimated for this study yields very similar marginal effects for education, potential labour market experience and location (rural/urban). This paper extends previous work on India in two ways. Regular and casual workers are considered separately as these workers have distinctly different wage determining processes.

[3] In their regression of wage premia on trade Goldberg and Pavcnik (2003) do not control for time-variant industry-specific factors other than sector-specific capital. If these are correlated with tariffs as well as wages the tariff coefficients could be biased though bias alone would not explain the sign changes. In this paper variables such as productivity, skill intensity and average enterprise size are used to control for such factors.

Estimates of the wage effects of these and other characteristics for all workers taken together mask potentially important differences between them. The other important distinction between this and earlier studies in India is that this paper focuses on the variation of wages across industries and investigates the determinants of this variation, particularly focussing on the trade reforms of the 1990s.

3. Trade liberalisation and the Indian economy

The key elements of India's pre-reform development model were rapid industrialisation with the prioritisation of capital goods over consumer goods, state control and regulation over the economy, and inward-orientation. Tentative attempts at reforming the Indian economy were made in the late 1980s especially with respect to industrial deregulation. However, the 1980s ended with a severe macroeconomic crisis in the winter of 1990-91 that necessitated a drastic stabilisation and structural reform programme (Kapila, 2001). The latter focussed primarily on trade and industrial policy reform but also encompassed taxation and financial reforms as well as institutional reforms relating to reform of labour, company, rent and land control laws and the establishment of adequate regulatory bodies (Ahluwalia, 2002; Srinivasan and Bhagwati, 1993).

India's major external sector reforms were in the following areas (Misra and Puri, 2001):

- the rationalisation and unification of the exchange rate in 1993 and the liberalisation of foreign exchange controls,
- the removal of restrictions on foreign investment,
- export promotion and the establishment of export processing zones,
- the elimination and tariffication of quantitative restrictions, and
- the reduction of tariff barriers.

<u>Trade policy</u>

The Indian tariff structure comprises a basic import duty (these are the statutory most favoured nation tariffs) and an auxiliary import duty on all imports. The tariff rates are mainly *ad valorem* with a few specific and composite rates. Additional import duties or countervailing duties (equivalent to the excise duty on like goods produced or manufactured in India) are also levied. A few goods are also subject to export taxes. The tariff structure is complicated by the presence of numerous exemptions such as general, end-use, specific user and preferential area exemptions. During the early 1990s there was substantial reductions in levels and dispersion of tariff rates as well as in the number of exemptions (Nouroz, 2001). Average tariff rates rose marginally during the mid-1990s as special customs duties were imposed after 1994. The special customs duties were eliminated in 1999 but a 10% customs surcharge was introduced on all imports, with some exemptions (Jain, 1999). During the 1990s on the whole, however, the peak and average tariff rates fell as did the dispersion between different tariff lines.

Table 1: Evolution of tariff rates (%)

	1983/84	1990/91	1991/92	1992/93	1993/94	1994/95	1995/96	1996/97	1997/98	1998/99	1999/2000
Maximum tariff [a]	135	355	150	110	85	65	50	52	45	40	40
Average unweighted tariff:											
Agriculture	73	113	..	..	43	..	27	26	26	30	29
Mining	57	100	..	..	70	..	30	26	25	29	27
Manufacturing [a]	103	126	..	..	73	..	42	40	36	41	40
Whole Economy [a]	98	125	..	..	71	..	41	39	35	40	40
Dispersion of tariff [b]	30	41	..	..	30	..	19	19	15	15	14
No. of tariff lines under quantitative restrictions [c]			4000 (80%)				< 1000				
NTB coverage ratio [d]							65.51		64.03	62.16	24.24

Source: Gulati (2000), Jain (1999; 1993); Kohli et al. (1983), Pandey (1998) and Rajan and Sen (2001).
Notes: These tariff rates include all auxiliary and special customs duties and customs surcharges where applicable. Only ad valorem rates are included. Year beginning 1 April. a\ These rates exclude some lines in the beverages manufacture sector that are subject to very high tariffs (e.g. the maximum rates were 193, 224 and 123% in 1983-84, 1993-94 and 1999-2000 respectively). b\ Measured by the standard deviation of the unweighted tariffs. c\ Based on the 6-digit Harmonised System code. d\ This is the percentage of commodities within a category that are affected by any form of a non-tariff barrier.

The extent of non-tariff barriers (NTBs) was high during the 1980s (Pandey, 1998) but during the late 1980s there was a move towards a tariff-based system for capital goods, intermediates and components. By the mid-1990s these goods were no longer subject to import licensing and could be freely imported on the OGL list (Kalirajan, 2001). Imports of consumer goods, however, remained virtually banned during the 1980s and had the highest NTBs even in the 1990s. Most of the NTBs on the mining[4] and manufacturing sectors were removed or decreased substantially in 1991. Though imports of most agricultural commodities are still on the 'canalised list'[5] or subject to licensing while exports are subject to minimum prices and quotas, many of the quantitative restrictions on agricultural trade were eased after 1994. The

[4] With the exception of crude oil and gold (Pandey, 1998).
[5] This consists of items that are imported or "channelised" through state agencies only such as agricultural commodities like grains, cereals, edible oils, oilseeds, sugar as well as non-agricultural commodities like petroleum products and fertilisers (Kalirajan, 2001).

coverage ratio for agriculture fell from 100% in the late 1980s to 60% in 1999-2000 (Pandey, 1998). Most agricultural and food imports (except beef and tallow) were removed from the prohibited list in 1997 (Kalirajan, 2001).

Industrial policy

The industrial regulatory policy until the mid-1980s was highly restrictive in terms of the decisions regarding capacity expansion, product mix and location decisions of firms. Industrial policy reforms were with respect to licensing, capacity expansion, small-scale sector regulations, the role of public sector enterprises, large firms and foreign investment (Kapila, 2001).

The Industries (Development and Regulation) Act, 1951, required the owner of any industrial undertaking to obtain a license (or permit) from the government in order to start production, produce a new product, expand existing capacity or enter a new market. Essentially this allowed the government control over private investment. Firms below a certain size of fixed investment were exempt from licensing requirements and firms classified in the small-scale sector were given additional privileges including the reservation of certain items for their exclusive production. There were additional restrictions on location and on investments made by large business houses (all new production by large business houses was to be confined to certain industries only). Schedules A and B of the IDRA defined state-dominated industries – the former listed 17 industries which were reserved for production by the public sector only while the latter listed industries that had some private sector participation but were predominantly state-led (Desai, 1992).

"Industrial licensing generated considerable red tape as well as strong political pressure; under the influence of both it underwent frequent tightening and relaxation" (Desai, 1992, pp. 112). Systematic deregulation began in earnest in the mid-1980s. In 1985 a system of "broad-banding" was introduced that allowed existing license-holders to diversify into a number of related industries without obtaining prior permission. By 1989 27 items were still subject to licensing. Further deregulation in the 1990s brought this down to 18 items in 1993 and finally to only 6 in 1999 (Misra and Puri, 2001; Sandesara, 1992). The minimum size of firms qualifying for small-scale sector concessions was raised for firms engaged in export. The restrictions on investment by large business houses were gradually relaxed in the late 1980s provided these houses generated sufficient export revenue or located in backward areas and the minimum asset limit was raised. In 1991 the distinction between business houses and other companies was eliminated and the rules pertaining to industrial location were also relaxed (Desai, 1992).

The reforms in 1991 also sought to reduce the role of the public sector by abolishing Schedule B and reducing the number of items reserved for the public sector alone, i.e. the Schedule A industries, from seventeen in 1983 to six in 1993 and finally to four in 1999. The aim was to limit public sector participation to the provision of infrastructural services though the reform and privatisation of existing enterprises has been very slow (Ahluwalia, 2002; Basu, 1993).

Foreign direct investment deregulation

In the pre-reform period foreign investment in Indian companies was, with some exceptions, limited to 40% and required prior approval from the government. The entire process was ad hoc, non-transparent and lengthy. The reforms during the 1990s encouraged the inflow of foreign capital by allowing automatic approval in selected areas, laying down rules for approval in other cases and simplifying and expediting the procedure. Automatic approval of up to 51% equity was granted in high technology and high priority industries such as the metallurgical, capital goods, electronics, food processing industries as well as services with a high export potential subject to maximum limit for each sector. This list of industries was gradually extended during the decade to include consumer goods and by 1999, except for a small ‘negative list’, all investment projects were to be given automatic approval. The negative list consisted of those industries that were still subject to industrial licensing and projects in excess of the 24% limit for areas reserved for the small-scale sector. The Foreign Investment Promotion Board was established to consider 100% foreign equity (Misra and Puri, 2001).

Thus, firms operating in the Indian market in the early 1980s faced barriers to entry due to government control over private investment through the licensing regulations, reservation of production for the public sector and lengthy and opaque procedures for approving foreign direct investment that was further subject to a maximum limit of 40% of equity. In addition there are barriers to exit due to labour market regulations - the Industrial Disputes Act prevents closure of units and lay-off of workers without prior government approval. Though these restrictions on entry were gradually eased from 1985 onwards those on exit were not brought under the liberalisation agenda until 2001 (Kalirajan, 2001).

Economic outcomes

The reforms have had a substantial impact on the Indian economy and generated high growth during the 1990s characterised by increasing exports and foreign investment. GDP had grown at 3% p.a. between 1950-51 and 1979-80, but the boom in the late 1980s (with 5.8% p.a. GDP growth) was fuelled by soaring fiscal deficits, monetisation of public debt, consequent inflation and balance of payments deficits (Kapila, 2001). After a sharp fall in 1991-92 following the crisis, the economy recovered and GDP grew by nearly 7% p.a. during the phase of rapid trade liberalisation over 1992-97 and by 6% p.a. between 1997-98 and 1999-2000. The policies of industrial deregulation and trade liberalisation triggered off strong export growth averaging 14% p.a. between 1992 and 1997 but decelerating to about 4% thereafter. Imports fell steeply till 1992 due to the imposition of strict controls to stabilise the economy but grew rapidly thereafter at about 15% p.a. in the first phase and then decelerating along with exports to about 7% p.a. The trade to GDP ratio was risen gradually from 14.8% to 20.6% between 1990 and 1999 (Ministry of Finance, various years).

Shifts in the composition of exports reflect the nature of the reforms undertaken. The share of agricultural exports in total exports rose as a result of the decanalisation and liberalisation of some commodities. Rice exports in particular increased to about 3% of total exports after private traders were allowed to engage in rice trade in 1994. India's share of the world rice market jumped from 9% to 12% in value terms making her the third largest exporter (Nielson, 2002). The share of manufactures in total exports has nearly doubled from 41% in 1990-91 to 78% in 1999-2000. The largest increases in this sector were in textiles and garments (reflecting India's comparative advantage in labour-intensive exports and the inflow of foreign investment (Rajan and Sen, 2001)), gems and jewellery (as a consequence of the export promotion measures introduced in the 1992-97 export-import policy), chemicals, and machinery (including computer software). With respect to imports, the share of capital goods in total imports nearly halved from 24% to 11% between 1990-91 and 1999-2000 while that of raw materials and intermediate goods rose (Ministry of Finance, various years).

This considerable structural change over the space of a decade is likely to have significant implications for the structure of the economy. The bulk of the adult male labour force in India is self-employed – about 58-60% of the labour force in rural and 35-39% in urban areas.[6] A

[6] Calculations for adult males aged 15-65 years are based on the current weekly status in the National Sample Survey employment survey data, 1983, 1993-94 and 1999-2000 (see Section 5 for details).

small proportion (about 3%) is unemployed. Of those in wage employment, regular wage or salaried workers comprised about 25% of the labour force while casual or contractual workers comprised about 22% of the labour force in 1983. The fall in the share of regular workers in the labour force to 22% accompanied by a rise in the share of casual workers to 24% by 1999 has raised concerns that the liberalisation of the economy led to increased casualisation. This paper focuses on the labour market outcomes of regular wage workers. Table 2 below outlines the trends in employment and real wages of regular workers in the economy (covering both the organised and unorganised sectors) by one-digit industrial classification.[7]

Table 2: Employment and real wages for regular workers by industry

	Employment share (%)			Real weekly wages [a]		
	1983	**1993**	**1999**	**1983**	**1993**	**1999**
Agriculture and allied sectors	11.10	5.07	5.67	53.24	81.83	125.24
Mining and quarrying	2.06	2.31	1.75	177.45	241.23	368.59
Light manufacturing	11.40	10.29	12.30	119.27	144.55	171.48
Heavy manufacturing	12.87	14.11	13.34	161.37	211.41	238.35
Utilities	2.22	2.39	2.31	180.16	259.47	431.95
Construction	1.74	1.66	1.83	138.04	189.55	227.55
Trade and hotels	7.73	8.65	12.34	89.45	110.50	169.94
Transport, storage and communication	11.59	10.78	11.73	150.22	191.16	249.84
Services	39.29	44.74	38.73	170.09	246.85	366.14
Economy	100	100	100	141.48	204.38	272.67

Source: National Sample Survey employment surveys, 1983, 1993-94, 1999-2000. a\ Nominal wages have been deflated to constant 1983 prices by the official consumer price indices (see Section 5 for details).

This dispersion in wages across industries and the impact of trade reforms on these inter-industry wage premia is explored in the following sections.

4. Methodology

This paper estimates inter-industry wage premia in the Indian labour market and examines the effect of the trade liberalisation during the 1990s on these premia. Following the standard labour economics literature wage regression models are estimated as augmented Mincerian earnings equations controlling for human capital, various working conditions and industry affiliation. Before the wage regression models are estimated the issue of potential selection

[7] Data on total earnings of workers in registered factories covered by the Annual Survey of Industries (i.e. the organised manufacturing sector) and deflated by the all-India CPI for industrial workers (see Section 5) indicates that the real weekly wage for these workers are much higher than those reported in Table 2: Rs. 168 and Rs. 294 in 1983 and Rs. 221 and Rs. 401 in 1999 for light and heavy manufacturing respectively.

bias is addressed.[8] The inter-industry wage premia are obtained from the wage regression model estimates following Krueger and Summers (1988) as deviations from an employment-weighted mean. These estimated industry wage premia are then regressed on trade variables and a variety of industry-specific characteristics in order to assess the role of trade in determining these premia.

4.1. Estimating wage regression models

The potential problem of selection bias is addressed here using the generalised framework popularised by Lee (1983). Consider the following two-stage model for selection and wage determination (suppressing the *i* subscripts for individuals):[9]

$$w_j = x_j'\beta_j + d_j'\delta_j + \mu_j \qquad j = 2, 3 \qquad (1)$$

$$y_s^* = z_s'\gamma_s - \eta_s \qquad s = 1,2,3 \qquad (2)$$

where w is the outcome variable (in this case, wages) for persons engaged in wage employment of two types – regular wage employment (j = 2) and casual wage employment (j = 3). As this bias is mediated through observed wages it is sufficient and more efficient to separate employment status into non-wage earners and different types of wage earners. The latent dependent variable, y_s^*, represents employment status of the individual - (1) non-participants in the labour market, self-employed and unemployed individuals, i.e. non-wage earners (2) regular wage employment and (3) casual wage employment. The outcomes, w (i.e. wages), are observed only if the person is in either form of wage employment (i.e. j = 2,3). The vector x_j and z_s are (Nx1) comprise exogenous explanatory variables, d_j are the (Kx1) industry dummy variables, s is a categorical variable signifying selection between the above 3 different alternatives, j is a categorical variable indicating regular or casual wage employment, μ_j and η_s are random error terms such that $E(\mu_s \mid x_j; z_s) = 0$ and $E(\eta_s \mid x_j; z_s) = 0$.

[8] The sample of individuals over which a wage function can be estimated is essentially truncated as data on wages as well as industry affiliation is reported only for those individuals in wage employment. If the selection of this sub-sample of individuals is random then an ordinary least squares procedure provides consistent and unbiased estimates of the coefficients. If this selection of individuals into wage employment is systematic (i.e. the error terms in the selection equation and the wage equation are correlated in some way) then ignoring the non-random nature of the sample would introduce a selectivity bias in the wage regression model's estimates.

[9] This sub-section is based on Lee (1983).

If the (η_s)s are assumed to be independent and identically distributed as Type I extreme value distributions then their difference (between different employment status) follows a logistic distribution. This gives rise to the conditional MultiNomial Logit (MNL) model and the probability that individual *i* experiences outcome *s* can then be expressed as:

$$P_s = \frac{\exp(z_s'\gamma_s)}{\sum_{j=1}^{M}\exp(z_j'\gamma_j)} \qquad s, j = 1,\ldots,3 \qquad (3)$$

The MNL model is identified only up to an additive vector so that adding the same vector to each of the coefficients (γ_s)s would generate the same probabilities. As a result one set of parameters (γ_s) must be selected as the base category and set to zero in order to overcome the indeterminacy inherent in the MNL model. Equation (3) then reduces to the following:

$$P_1 = \frac{1}{1+\sum_{j=2}^{M}\exp(z_j'\gamma_j)}; \quad \text{and} \quad P_s = \frac{\exp(z_j'\gamma_s)}{1+\sum_{j=2}^{M}\exp(z_j'\gamma_j)} \qquad \gamma_1 = 0;\ s = 2,3;\ j = 2,3 \qquad (4)$$

In this paper outcome 1, i.e. non-participants, self-employed and unemployed persons, is taken as the base category and the other two sets are estimated relative to this category. In order to identify the parameters of the wage equations a set of variables that influence employment status between the alternative outcomes but not wage itself must be included as regressors in the selection equation. These are usually variables that capture exogenous household non-labour income and/or family background (e.g. parent's education level). In the absence of data on income from different sources (other than wages) and on family background alternative identifying variables have been used in this paper (see Section 6.1 below).

Consistent estimates of the parameters (β_1 and β_2)s of equation (1) can be obtained by replacing the disturbance terms μ_1 and μ_2 in equation (1) by their conditional expected value obtained from the MNL estimation (equation 4). This selection bias correction term, λ_j, is similar to the inverse of the "Mills ratio":

$$\lambda_j = \frac{\phi(J_j(z_j^{'}\gamma_j))}{F_j(z_j^{'}\gamma_j)} = \frac{\phi((J_j(z_j^{'}\gamma_j)))}{F_j(z_j^{'}\gamma_j)} \qquad j = 2, 3 \qquad (5)$$

where $\phi(.)$ and $\Phi(.)$ represent the normal density and distribution respectively, *J(.)* represents the 'normits' or the standardised z-scores for each observation, i.e. $J(.) = \Phi^{-1}(F(.))$. It follows that $J_j(z_j^{'}\gamma_j) = \Phi^{-1}(P_j)$ where P_j is the probability of being in outcome j (j = 2, 3) and the selection bias correction term can be re-written as:

$$\lambda_j = \frac{\phi(\Phi^{-1}(P_j))}{P_j} \qquad j = 2, 3 \qquad (6)$$

Thus, using the predicted probabilities from the reduced form MNL model the selection bias correction term, λ_j, can be constructed for each individual for outcomes 2 and 3, i.e. regular and casual wage employment, and included in the corresponding wage equations to control for potential selection bias.

A standard semi-logarithmic Mincerian specification can then be used to estimate the wage equations (Mincer, 1970):

$$\ln(y_j) = x_j{}'\beta_j + d_j{}'\delta_j - \beta_j^* \hat{\lambda}_j + \upsilon_j \qquad j = 2,3 \qquad (7)$$

where subscript j = 2,3 refers to regular and casual workers respectively; $\ln(y_j)$ is the natural log of wages, $\beta_j^* = \rho_j \sigma_{\mu_j}$ the coefficient on the selection bias correction term in the wage equations; ρ_j the coefficient of correlation between the error terms in the wage equation and the selection equation (the direction of bias is determined by this correlation term); and υ_j the error term for each of the wage equations.

This two-step procedure controls for the underlying process by which the set of observations actually observed are generated. It ensures that the OLS estimates of the coefficients from the wage equations are consistent. The sampling distribution for the estimates can be obtained by using a modification to the formula suggested in Trost and Lee (1984) or by bootstrapping as suggested by Bourguignon et al. (2001). Each of the wage regression models in this paper has been bootstrapped using 1000 replications.

4.2. Estimating inter-industry wage premia

In a methodology first introduced by Krueger and Summers (1988) and followed in empirical studies of inter-industry wage structures this paper extracts inter-industry wage premia from the wage regression models estimated above. This paper differs slightly from the standard approach in that all industry dummies are included in the wage regression models in the absence of a constant. This allows the model to be interpreted as a fixed effects model where the industry effects capture omitted factors. The coefficients on the industry dummies are then normalised as deviations from a employment-weighted mean differential as follows:

$$\hat{\delta}^* = (I - s')\hat{\delta} \tag{8}$$

where $\hat{\delta}^*$ is the column vector of industry wage premia, I is a identity matrix, $\hat{\delta}$ is the column vector of industry coefficients estimated from the wage equation (7), and s is a matrix of industry employment weights with each element $s_k = n_k \Big/ \sum_{k=1}^{K} n_k$ where n_k is the share of regular workers in industry k and all matrices have (Kx1) dimensions for k=1,..K industries. The adjusted variance-covariance matrix $V(\hat{\delta}^*)$ is computed as suggested by Haisken-DeNew and Schmidt (1997)[10] and can be expressed as follows:

$$V(\hat{\delta}^*) = (I - s')V(\hat{\delta})(I - s)' \tag{9}$$

The resulting wage premia represent the difference in the wage received by a worker in industry j to the average worker across all industries in the economy. The overall variability in wages across industries can be approximated by the average employment-weighted adjusted standard deviation of the wage premia computed as follows:

[10] Krueger and Summers (1988) suggest using the standard errors of the estimated industry coefficients δ 's (and that of the constant for the omitted industry) from the wage regression. Haisken-DeNew and Schmidt (1997) show that this overstates the standard error and that the degree of bias is sensitive to the choice of the omitted industry. Also, the summary variable of dispersion - the employment-weighted adjusted standard deviation of the wage premia - is likely to be underestimated though this is not as sensitive to the choice of base industry in larger samples.

$$SD(\hat{\delta}^*) = \sqrt{s'(Diag(\hat{\delta}^*))\hat{\delta}^* - s'Col(V(\hat{\delta}^*))} \tag{10}$$

where *Diag(.)* transforms the (Kx1) column vector into a (KxK) square matrix with the diagonal elements given by the column vector and *Col(.)* denotes the column vector formed by the diagonal elements of the matrix.

4.3. Estimating the determinants of inter-industry wage premia

In order to explore the impact of trade liberalisation on the inter-industry wage structure for regular workers the estimated wage premia are regressed on tariffs (t_j). Various controls for industry-specific characteristics (z_j) are also included in a more general specification. The observations are pooled across the three years so as to obtain more efficient estimates. Since the dependent variable is estimated from the wage equation it raises concerns that the coefficients in the wage-trade regression models might have large variances that could possibly differ across industries depending on the variance of the estimated industry coefficients. A weighted least squares (WLS) regression model is estimated. The weights used are the inverse of the variances of the estimated wage premia from the wage equations implying that sectors with larger variances are given lower weight in the estimation.

$$wp_k(\theta_k) = (\theta_k)t_k'\alpha + (\theta_k)z_k'\beta + (\theta_k)\xi_k \qquad k=1,\ldots,\text{K industries} \tag{11}$$

where the weights, $\theta_k = \sqrt{1/\text{var}_k}$, are the inverse of the variances of the estimated wage premia from the wage equations at time *t* so that sectors with larger variance are given lower weight (time subscripts suppressed).

5. A description of the data and variables used

This paper exploits three national employment surveys conducted by the National Sample Survey Organisation (NSSO) for 1983, 1993-94 and 1999-2000. The first survey can be interpreted as providing insights into the structure of Indian labour markets prior to liberalisation while the latter two provide the basis for delineating a portrait of these structures after the radical trade liberalisation process.

Employment surveys

The key years of this analysis are the years that the large-scale employment surveys were undertaken - January-December 1983, July 1993–June 1994 and July 1999–June 2000 (referred to as 1983, 1993 and 1999 in this paper).[11] These surveys provide comprehensive national coverage and provide a wealth of information on numerous socio-economic issues at the household and individual level. This survey period is split into four sub-periods of three months duration each, corresponding approximately to the four climatic seasons in 1993-94 and 1999-2000 and with the four agricultural seasons in 1983 (National Sample Survey Organisation, 1987; 1997; 2001). These surveys cover all workers but do not have information on whether the worker is employed in the organised or unorganised sector (except for the last year).

Wages

Nominal weekly wages include payment in cash and kind. Some observations (about 1-2% in the three years) had to be dropped from the sample used in the wage regression models as there were missing observations on wages, hours worked and industry affiliation.[12] The wage distribution was then trimmed by 0.1% at the top and bottom tails.[13] These nominal wages were deflated to 1983 prices using official state-level monthly consumer price indices (base year 1960-61) for agricultural labourers (CPIAL) for rural wages and industrial workers (CPIIW) for urban wages (Labour Bureau, various years).[14] The employment surveys have data on the intensity of work – i.e. 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) – for each day of the week. Using this information and assuming a 48 hour week, the number of hours worked and the hourly real wage was constructed.

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[11] The employment survey for 1987-88 could not be used as over 76% of observations on rural wages for persons participating in wage employment are missing.

[12] It is assumed that the excluded observations are random as the mean observable characteristics of the workers excluded do not differ significantly from those retained in the sample though this does not take possible differences in unobservables into account.

[13] This is necessarily an ad hoc measure: some researchers prefer to trim the wage distribution using specific values (Krueger and Summers, 1988) while others prefer to trim the distribution at the tails (Dickerson et al., 2001) as adopted here.

[14] Deaton and Tarozzi (1999) and Özler et al. (1996) criticise these indices as the weighting diagrams have remained unchanged for many years and these indices do not take into account state-level rural-urban cost of living differentials. However, alternate indices cannot be used as Deaton and Tarozzi (1999) do not compute price indices from the NSS survey data for 1983 while Özler et al. (1996) do not publish corrected official indices for 1999.

Variables influencing wages

The sample is restricted to prime age males aged between 15 and 65. Age splines at ten-year intervals were included to capture age effects as a proxy for labour force experience. Individuals were divided into three mutually exclusive categories using current weekly status: (1) non-participants in the labour market, self-employed and unemployed individuals, i.e. those without wages (2) casual wage employment and (3) regular wage employment. There is information on the highest level of schooling completed (but not on the number of years of schooling) so dummy variables corresponding to the following education variables were constructed: primary school, middle school, secondary school and graduate and above. The reference category is individuals who are illiterate or have less than two years of formal or informal schooling. Dummy variables for the quarter in which the households were interviewed are constructed to counter possible seasonality effects arising out of the fact that households were interviewed at different points in time. These quarterly dummies were also interacted with the dummy variable for the rural sector.

Industry affiliation

The variable for industry affiliation was constructed based on the individual's current weekly industrial classification (3-digit National Industrial Classification, NIC). A concordance table was constructed between the three different revisions of this code (Central Statistical Organisation, 1987a) and was mapped into the Input-Output tables for 1983-84 and 1993-94. The tables for both years are constructed for 115 production industries and are strictly comparable over time and can be matched with the industry affiliation for persons reported in the NSS employment surveys. In order to ensure adequate observations in each industry the original 115 industries from the Input-Output were aggregated so as to obtain 54 industries for regular wage workers.

Tariff data

This paper captures the effect of trade policy reform using simple averages of tariffs for 1983-84, 1993-94 and 1999-2000 (Jain, 1993; 1999; Kohli et al., 1983). This includes the basic customs tariff for all three years as well as the auxiliary tariff in 1983-84[15] and the surcharge on basic customs tariff in 1999-2000.[16] The additional duty is not included because it is a

[15] In 1993-94 the auxiliary duty was merged into the basic customs tariff.

[16] Specific and composite rates could not be quantified and have not been included. In the case of composite rates, the ad valorem rate is included. There were few such rates as most tariffs are ad valorem.

countervailing duty for excise duties on domestic production (Pandey, 1998). The tariff structure is complicated by the presence of numerous exemptions that may be complete or partial and may apply to either the entire tariff line or only to certain items within the tariff line. Those exemptions that are quantifiable and are applicable to all goods within a tariff line are taken into account in this analysis. A concordance table was constructed between the Brussels Tariff Nomenclature (BTN) used in the 1983 tariff schedule and the Harmonised System (HS) used in the other two years.[17] The tariff data was then mapped into sectors (Central Statistical Organisation, 1990). Tariff data are available for the tradable sectors, i.e. agricultural, mining and manufacturing industries.

Industry-specific variables

These variables are constructed for the second-stage wage-trade regression models for manufacturing industries only. Two dummy variables are constructed to capture the incidence of industrial policy. The industrial license dummy is coded 1 for industries that have to obtain a license (or permit) from the government in order to start production, expand capacity or enter a new market and 0 otherwise. In 1983 all manufacturing industries were subject to licensing requirements. The public sector dummy is coded 1 if the industry is a Schedule A industry, i.e. reserved for the public sector and 0 otherwise. Trade flow variables for 1983 and 1993 were constructed from the input-output tables for these two years. In the absence of a more recent input-output table than 1993-94, trade data for 1999 were constructed from data on imports and exports values at the 6-digit HS level (DGCIS, various years).[18]

Variables controlling for industry characteristics such as the share of female, casual and unskilled workers in the industry labour force are constructed from the NSS survey data.[19] Data on other industry-specific variables such as labour productivity, average establishment size and union density are available only for the organised manufacturing sector and are obtained from the Annual Survey of Industries (Central Statistical Organisation, 1987b; 1997; www.indiastat.com, 2003a) and other secondary sources.[20] The capital-labour ratio is used to

[17] This was done via the concordance from HS to SITC-Revision 2 and from this to BTN (United Nations, 1975). This is not a perfect concordance as the BTN is far more aggregated than the HS and the same BTN heading often corresponds to more than one HS heading. In addition, the Indian tariff schedule combines some BTN headings together. The assumption is made that all goods within a tariff heading are subject to the same tariff.

[18] In cases where a particular HS line was common to two industries 8-digit HS codes and textual descriptions were used to assign the trade value to one of the industries.

[19] Skilled workers are defined as those who have completed secondary and/or graduate education.

[20] This data is for the financial year running from April-March that does not exactly match the timing of the NSS employment surveys. The industrial classification (3-digit NIC-1998) for the 1999 ASI data does not permit the

proxy labour productivity. The fixed capital stock in each industry is deflated to constant 1983 prices by the wholesale price index (WPI) for machinery, transport equipment and construction materials (i.e. products of the non-metallic minerals industry) (www.indiastat.com, 2003b).[21] Labour is measured as the total number of employees in each industry.[22] The average establishment size is computed as the number of employees per factory. Union density is constructed from data on the membership of trade unions deflated by the number of employees in the organised sector of each industry (Labour Bureau, various years).[23] States are classified into three groups – those that have passed pro-worker or anti-workers amendments or are neutral, i.e. did not experience any amendment activity - and the share of employment in total industry employment in the these state groups are included in order to capture a regional dimension at the industry level.[24]

6. Empirical Results

As mentioned earlier, the key years in this paper are 1983, 1993 and 1999. India's trade reforms were implemented rapidly between 1991 and 1994 and, after a short lull, in the late 1990s. Hence, the first survey provides a picture of the pre-reform structure of the Indian labour market while the latter two capture the post-liberalisation period. The empirical analysis reported in this chapter is restricted to prime-aged adult males.

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same mapping into sectors as the previous years. In four cases the total value of the variable in question is disaggregated into sectors based on the sectoral shares of that variable in 1993.

[21] Hasan et al. (2003) use a similar deflator.

[22] The data on the number of employees is essentially the average daily employment in the industry as it is constructed by dividing the total man-days by the number of days worked in the year.

[23] There are two sources for constructing the union density measure – one, as in this paper, from data based on trade union returns and ASI organised sector employment data, and two, from the NSS surveys. The first source suffers from problems of underreporting in the trade union data as not all unions are legally required to submit annual statutory returns and many state governments do not publish data on registered trade unions. The most recent union data available is for 1998. In the second case, though the 1993 and 1999 NSS surveys report union membership data this variable cannot be used in either the wage-trade auxiliary equations or the wage equations because nearly half the observations on regular workers have missing data. Also, there is no corresponding variable in the 1983 survey. Comparisons with the ILO estimates of union membership as a share of formal sector wage earners for 1980 and 1991 reveal that the union density estimated using the first method is more appropriate and that estimated using the second method is vastly overstated (ILO, 2002).

[24] Group one comprises the pro-worker states of Gujarat, Maharastra, Orissa and West Bengal; group two comprises the anti-workers states of Andhra Pradesh, Karnataka, Kerala, Madhya Pradesh, Rajasthan and Tamil Nadu; and group three comprises neutral states of Assam, Bihar, Haryana, Punjab and Uttar Pradesh. The Besley and Burgess (2002) dataset includes Jammu and Kashmir (neutral) but not Himachal Pradesh and Tripura. The mode of the labour regulation variable is six. The two omitted states are classified as neutral in the absence of information regarding labour law amendments.

6.1. Reduced form multinomial model

In the first step a reduced form multinomial logit (MNL) model estimates the probability of being in a particular state out of three unordered alternatives: (1) non-participation in the labour market, unemployment, or self-employment; (2) regular wage employment; or (3) casual wage employment. In order to resolve the indeterminacy in the MNL model the first outcome, i.e. non-participation, self-employment and unemployment in the labour market, is set as the base. The explanatory variables used in the multinomial logit are: worker characteristics such as age, level of education and marital status, controls for location (rural/urban and state of residence) and seasonality effects (proxied by the timing of the interview for the survey), and variables capturing social exclusion operating through caste and religion.[25] The parameters of the wage equations are identified using variables that capture household structure captured by variables on household size and dependency measures – three dummy variables for whether the household has one child, two children or three or more children aged 0-4 years (the omitted category is not having any children aged 0-4 years) and the number of persons aged more than 65 years in the household.[26]

It must be stressed that this estimation is not an attempt at modelling participation in the labour market but one designed to obtain the necessary tools to control for potential selection bias in the wage regression models. The results for the multinomial model for selection into both types of wage employment – casual and regular – are not reported here. The majority of the effects estimated are plausible and are significant at the 1% level or better. Individuals who are educated, married with a large number of children and reside in urban areas are more likely to be in regular wage employment. The direction of effect of most of the variables

[25] Social exclusion can be thought of as the process by which certain groups are continuously marginalized or excluded in society (Das, 2003). Nayak (1994) conceptualises the problem of social exclusion as being "one of lack of entitlement of economic and social power amongst a large section of the population" where "the notion of entitlement refers to the actual or effective empowerment of a person to trade his original endowment of labour power and other factor incomes for food and other basic necessities" (pp.2). There are many dimensions to exclusion; this paper considers exclusion from the labour market due to caste and religion. Other sources of exclusion such as gender and casualisation of work are irrelevant in this study as the sample is restricted to adult males in regular wage or salaried employment.

[26] As the choice of identifying variables is necessarily ad hoc the MNL model was estimated for different specifications of identifying variables and was found to be robust to the choice of household structure variables. The mean of the selection bias correction term was in a tight range of –0.001 to +0.013 around the mean estimated by the model reported here (1.149, 1.26 and 1.174 for the three years). The choice of the identifying variables did not seem to affect the size and sign of the coefficients in the wage equations and the coefficient on the correction term itself was insignificantly different across specifications.

remained stable across all three years with a few exceptions, mostly for the state effects in 1993.

6.2. Wage regression models

A standard semi-logarithmic Mincerian specification, augmented by a variety of controls, is used to estimate the wage equations (Mincer, 1970) for regular wage workers for the three years after correcting for potential selection bias. The dependent variable is the natural log of real hourly wages. The explanatory variables used in the wage equations, in common with the MNL model, are: worker characteristics such as age, level of education and marital status, caste and religion, controls for location (rural/urban and state of residence) and seasonality effects. In addition, the key variables of interest – industry affiliation of the worker – have been included in the wage equations. The summary statistics of the variables used in the wage equations are given in Table A1 in the Appendix.

Table 3 reports the results of the wage equations estimated for 1983, 1993 and 1999. All industry dummies are included but not reported in the regressions below as they are reported in Table A2 in the Appendix as deviations from an employment-weighted mean differential. The explanatory power of the variables in all three years is quite high though the fits appear poorer in the second year. The standard error of the estimate quantifies the deviation of data points around the regression plane and has increased by about 10 percentage points between 1983 and 1999.

Table 3: Wage regression models for regular wage workers

Dependent variable: Natural log of real hourly wages

	1983	**1993**	**1999**
Individual characteristics:			
Age: 15-25 years	0.0131***	0.0104***	0.0135***
	(0.0017)	(0.0024)	(0.0022)
Age: 25-35 years	0.0174***	0.0162***	0.0200***
	(0.0009)	(0.0012)	(0.0013)
Age: 35-45 years	0.0110***	0.0159***	0.0155***
	(0.0008)	(0.0010)	(0.0010)
Age: 45-55 years	0.0055***	0.0134***	0.0166***
	(0.0013)	(0.0014)	(0.0014)
Age: 55-65 years	-0.0287***	-0.0303***	-0.0251***
	(0.0031)	(0.0052)	(0.0046)

	1983	1993	1999
Married	0.0660***	0.0772***	0.0867***
	(0.0070)	(0.0094)	(0.0098)
Education:			
Completed primary school	0.0669***	0.0426***	0.0479***
	(0.0073)	(0.0098)	(0.0111)
Completed middle school	0.1381***	0.0939***	0.1074***
	(0.0084)	(0.0121)	(0.0114)
Completed secondary school	0.3503***	0.2652***	0.2942***
	(0.0110)	(0.0160)	(0.0140)
Completed graduate school	0.6115***	0.5338***	0.6003***
	(0.0153)	(0.0227)	(0.0188)
Social exclusion:			
Member of scheduled caste or tribe	-0.0449***	-0.0385***	-0.0320***
	(0.0055)	(0.0073)	(0.0072)
Muslim	-0.0129*	-0.0293***	-0.0417***
	(0.0075)	(0.0096)	(0.0087)
Seasonality:			
Household interviewed in (season):			
2nd quarter	-0.0234***	0.0039	0.0044
	(0.0073)	(0.0080)	(0.0094)
3rd quarter	-0.0311***	-0.0269***	-0.0061
	(0.0075)	(0.0093)	(0.0095)
4th quarter	-0.0440***	-0.0066	-0.0051
	(0.0076)	(0.0086)	(0.0093)
Rural * season2	-0.0091	0.0121	0.0147
	(0.0121)	(0.0155)	(0.0169)
Rural * season3	-0.0141	0.0680***	0.0055
	(0.0127)	(0.0168)	(0.0172)
Rural * season4	0.0334***	0.0260	-0.0243
	(0.0128)	(0.0160)	(0.0171)
Location:			
Residence in rural areas	-0.1367***	-0.0365*	-0.0176
	(0.0142)	(0.0198)	(0.0175)
State of residence:			
Andhra Pradesh	-0.0926***	-0.1364***	-0.2435***
	(0.0102)	(0.0124)	(0.0144)
Assam	0.0055	0.0340**	-0.0495***
	(0.0133)	(0.0162)	(0.0176)
Bihar	-0.0346***	0.0646***	-0.0277
	(0.0118)	(0.0189)	(0.0173)
Gujarat	0.0271**	-0.0065	-0.0796***
	(0.0111)	(0.0135)	(0.0144)
Haryana	0.0837***	-0.0498**	0.0645***
	(0.0170)	(0.0207)	(0.0225)
Himachal Pradesh	0.1603***	0.0279	-0.0025
	(0.0199)	(0.0207)	(0.0184)
Karnataka	-0.0338***	-0.0524***	-0.1567***
	(0.0126)	(0.0142)	(0.0152)
Kerala	0.0332**	-0.0364**	-0.0583***
	(0.0131)	(0.0176)	(0.0174)
Madhya Pradesh	-0.1293***	-0.1497***	-0.2634***
	(0.0099)	(0.0127)	(0.0151)

	1983	1993	1999
Maharashtra	0.0376***	0.0251**	-0.0518***
	(0.0094)	(0.0106)	(0.0135)
Orissa	-0.0420***	0.1689***	0.1033***
	(0.0139)	(0.0172)	(0.0211)
Punjab	0.0831***	0.0468***	-0.0091
	(0.0115)	(0.0138)	(0.0159)
Rajasthan	-0.0041	-0.0168	-0.0696***
	(0.0124)	(0.0161)	(0.0153)
Tamil Nadu	-0.1079***	-0.0362***	-0.0891***
	(0.0106)	(0.0125)	(0.0145)
Tripura	0.0492**	-0.0633***	-0.1089***
	(0.0231)	(0.0238)	(0.0221)
Uttar Pradesh	-0.0869***	-0.0207	-0.1573***
	(0.0104)	(0.0137)	(0.0137)
Selection bias correction term	-0.0243	-0.1164***	-0.1125***
	(0.0174)	(0.0266)	(0.0226)
Selection effect	0.0279	0.1310***	0.1321***
	(0.0230)	(0.0337)	(0.0311)
Number of observations	27,356	26,387	27,295
R2	0.5521	0.4852	0.5345
Standard error of estimate	0.3464	0.4210	0.4461

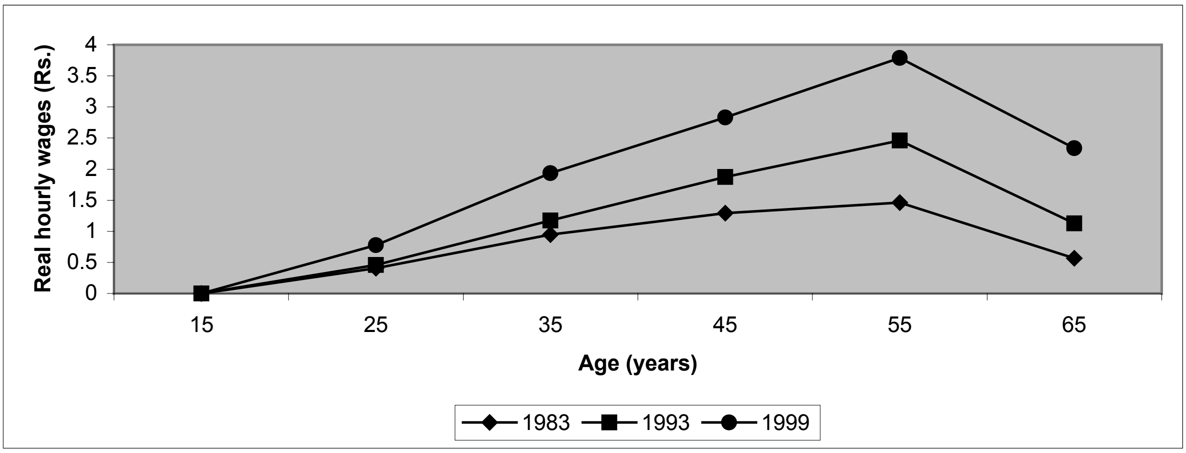
Notes: 1/ Standard errors in parentheses 2/ * significant at 10%; ** significant at 5%; *** significant at 1%. 3/ Industry coefficients not reported as the inter-industry wage premia are reported in Table A2 in the Appendix. 4/ The estimated coefficients on the age splines are not cumulative. 5/ The above regressions were estimated without a constant and with all industry dummies. The adjusted R^2 reported here is computed as the squared correlation between the actual and predicted values of the dependent variable. 6/ The selection effect is computed as the average selection bias correction term times its estimated coefficient. A crude estimate of the standard error of the selection effect is obtained as follows: the square of the average selection bias correction term times the standard error of its estimated coefficient.

Individual characteristics

Age serves as a proxy for labour market experience as the employment surveys do not report data on actual labour market experience and there is insufficient information to construct a potential labour market experience variable. For adult males workers age is likely to be highly correlated with labour market experience. The standard quadratic form for age is not used as this did not fit the data well and following Murphy and Welch (1990) age splines at ten-year intervals were included instead.[27] Figure 1 below reveals that the age-earnings profile display a positive relationship between age and real hourly wages and the general shape in accordance by human capital theory and previous empirical research (Murphy and Welch, 1990).

[27] Ten year intervals were chosen in order to maintain a balance between tighter splines and comparability between the three years and with other work on wage determination for casual wage workers.

Figure 1: Age-earnings profile, 1983-1999



The age-earnings profiles in the three years are concave indicating that wages increase at a declining rate till they reach a peak and start falling.[28] The returns to age rise steeply initially for the 25-35 age group (more so in the 1990s), then rise at a declining rate to peak at the 45-55 age group and finally decline for the 55-65 age group. The age-earnings profile has clearly shifted up during these three years - a Wald test of the coefficients on the age splines rejects the null hypothesis of no movement between each pair of years.[29] The age-earnings profile is steeper during the 1990s, particularly in 1999, with wages rising much faster as workers age relative to 1983. This combined with the rising standard error of estimates in the wage equations indicate an increase in the returns to unobservable skills that are possibly related to the liberalisation process.

<u>Education</u>

The omitted category for the education dummy variables is those who are illiterate or have less than two years of any type of education. The marginal wage effects of education for regular workers are significantly positive and monotonically increasing in education level – a regular worker educated till primary school earned wages about 7% higher than one with no education while a graduate earned as much as 61% higher wages in 1983. This impact at every level of education fell sharply between 1983 and 1993. The general trend appears to be that skilled workers with regular jobs fared badly in 1993 relative to illiterate, just literate workers and workers educated till primary school and the gap between the lowest and highest

[28] Since these are cross-section data the age-earnings profile refers to the wage received by different workers at different age groups. They do not trace the earnings of an individual worker over time. Comparisons of returns to age over time are for cohorts of workers falling in the same age groups in the three years.

[29] The χ^2 statistics (5 degrees of freedom) are 85.85 between 1983 and 1993, 20.75 between 1993 and 1999 and 186.16 between 1983 and 1999.

education levels – primary and graduate school – fell in this year. By 1999 the marginal effects of education had risen once more (though this rise was significant only for graduates) and this gap widened again.

Direct comparisons with earlier studies on the Indian labour market are not possible as these estimate different wage regression models. Duraisamy (2002) regresses the log of daily nominal wages on a smaller set of explanatory variables for all workers, regular and casual, without controlling for possible selection bias when analysing trends between 1983 and 1993. Kingdon and Unni (2001) also analyse the returns to education in 1987-88 but for urban areas in just two states, Tamil Nadu and Madhya Pradesh,. This paper controls for potential selection bias, allows for different wage determination processes for regular and casual workers and includes a number of additional determinants, mainly the industry effects. Duraisamy (2002) finds that the wage effects of all levels of education above primary school fall between 1983 and 1993 while that of primary education rise.[30] Both studies find that the education effects are positive and monotonically increasing for male workers.

How does this tie in with trade theory predictions on wage inequality? The Stolper-Samuelson theorem predicts a reduction in the skill premium and narrowing wage inequality with greater trade openness in developing countries like India that are relatively abundant in unskilled labour, assuming that Indian exports are more labour intensive relative to imports.[31] The results in Table 3 above suggest that the skill premium for regular workers fell immediately following the trade reforms but rose during the 1990s especially for graduates. These results are the opposite of those predicted by the Heckscher-Ohlin framework for an unskilled labour abundant country like India. Though it is not possible to attribute these changes to the effects of trade liberalisation during this period without formal analysis, it is likely that the dramatic turn-around of policy would have had a discernable impact.

The empirical evidence on trade liberalisation and changes in the skill premium in developing countries is mixed with the skill premium rising in some countries after reform and falling in others (see Wood (1997)). A possible explanation is that trade may be Skill-Enhancing-Trade or "SET" (i.e. trade induces developing countries to adapt modern skill-intensive technology

[30] A comparable specification of the wage equation estimated here replicates Duraisamy's results.
[31] The Heckscher-Ohlin framework is consistent with the presence of inter-industry wage premia provided these reflect non-pecuniary differences across industries that are fixed over time.

from developed countries). Since capital and skill are often complementary, the relative demand for skilled workers would rise (Robbins, 1996). It is possible this was the case in India – the share of basic and capital goods in India's imports rose from 43% to 50% between 1983 and 1993 and further to 60% in 1999.[32] Another possible explanation is that, as Harrison and Hanson (1999) find in Mexico, wage inequality between skilled and unskilled workers increased after the trade reform as unskilled labour-intensive sectors were highly protected by tariff barriers. After reform, the fall in these barriers worsened the situation of unskilled workers and contributed to widening wage inequality. In the Indian case as well the data suggests that manufacturing industries with the highest tariffs also had the largest share of unskilled workers. In addition the largest tariff reductions between 1983 and 1999 were in industries with the highest initial tariffs (see Section 6.4.1. below and Table A3 in the Appendix).

On the other hand, it is possible that the contraction of the skill premium for regular workers in 1993 was the effect of the trade reforms (and in line with the Heckscher-Ohlin predictions) while the gradual rise in 1999 reflected a trend, such as skill-biased technological change.

Social exclusion

It is hypothesised that socially disadvantaged groups would not only be excluded from the regular wage jobs (this is supported by the MNL model) (Das, 2003) but would also receive lower wages either due to lack of promotional opportunities or because of crowding into certain occupations within an industry (Nayak, 1994). Table 3 reveals that belonging to a scheduled caste or tribe or being Muslim significantly decreases the wage received in all three years. Kingdon and Unni (2001) find no significant effect of caste or religion in their study but while this may be true for urban areas in the two states they examined it might not be case for India as a whole. The disadvantage faced by Muslims in the labour market has increased over time. These negative wage effects could be due to correlation with unobservables, such as post-employment discrimination, or omitted variables such as occupation. Occupation variables have not been included in these wage regression models as the classification is very similar to the industrial classification. Traditionally individuals belonging to scheduled castes or tribes have been associated with low-wage occupations. The survey data reveals that the

[32] Basic and capital goods industries include the manufacturing of petroleum products, heavy chemicals, basic metals, machinery and transport equipment following the official classification used for constructing the index of industrial production (EPWRF, 2003).

largest proportion (and greater than the economy average) of scheduled caste workers are engaged in agricultural and allied occupations while Muslims are engaged in production, construction and transport work. Both these occupations paid less than the average wage in all three years.

Geographic and seasonal effects

Residing in rural areas significantly reduces the wage received by regular workers but this disadvantage fell over time and was no longer significant in the 1990s. The seasonal (interacted with settlement type) effects are jointly significant[33] - working in any quarter other than the first quarter of the year tends to reduce the wage received, though this is less pronounced or even reversed for rural areas. The state effects are largely significant[34] indicating the presence of either geographic, language or ethnic barriers to mobility or different institutional arrangements for wage-setting. Transforming these into deviations from an employment-weighted mean differential reveals that the inter-state dispersion in wages[35] increased from 7% to 10% between 1983 and 1999. Though there is inter-state migration of workers (about 4% of all workers and 12% of regular workers reported their current residence at the time of the NSS survey as different from their previous residence in another state in 1999)[36] the persistence of the inter-regional wage differentials as estimated above indicate the presence of constraints on inter-state mobility. There is some evidence that labour markets function differently across states (Aghion et al., 2003; Besley and Burgess, 2002).[37]

Selection effects

The selection effect is computed as the coefficient on the selection bias correction term times its mean for the nominated outcome, i.e. regular wage employment.[38] The selection effect is not significant for regular workers in 1983 but by the 1990s individuals selected into regular wage employment tend to earn about 13% higher wages than a person randomly selected from the population. Kingdon and Unni (2001) also find a significant positive selection effect for

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[33] The χ^2 statistics (6 degrees of freedom) are 60.46, 21.58 and 13.22 for the three years respectively.

[34] The χ^2 statistics (16 degrees of freedom) are 1000.79, 758.85 and 908.29 for the three years respectively.

[35] As measured by the adjusted employment-weighted standard deviation of the state premia.

[36] The inter-state migration flows for all workers and regular male workers were similar in 1983 at 3% and 11% respectively. There is no information on migration in the 1993 NSS survey.

[37] Hasan et al. (2003) also explore the differential impact of trade liberalisation across industries and states using data on manufacturing industries in the organised sector. Concerns of adequate observations in each category prevent the analysis of this question using interactions between state and industry dummy variables.

[38] It should be noted that the coefficient of the selectivity bias correction term in equation (7) has a negative sign.

all workers in their state samples. It is possible that individuals with unobservable characteristics such as better ability, motivation, etc. are absorbed into regular wage employment.

6.3. Inter-industry wage premia

The industry coefficients estimated in the wage regression models above are all significant at the 1% level. These are transformed into deviations from the employment-weighted mean differential as described in Section 4.3. and are reported in Table A2 in the Appendix. The inter-industry wage premia are large and the range between the wage premia in the lowest and highest wage industries has widened between 1983 and 1999 from 0.6380 to 0.8811 (in absolute values). The wage premia range from a maximum of 37% in the fuels mining sector to a minimum of -50% in the plantation crops sector in 1999 for regular workers. The high wage sectors for regular workers are services such as banking and insurance; heavy manufacturing such as petroleum, chemicals, metal, machinery and transport industries; mining and fuel extraction; and utilities. Low wage sectors for regular workers comprise of agricultural and allied sectors; light manufacturing such as foodstuffs, beverages and tobacco and textiles; wholesale and retail trade; hotels and restaurants; and services such as legal, business, personal and community services. Industry wage premia and the share in imports are positively correlated for all tradable as well as manufacturing industries.

High wage sectors tend to have a higher share of skilled workers and a lower share of casual workers and for manufacturing sectors, higher capital-labour ratios. These results are consistent with the findings of studies on other countries where capital and/or skill-intensive industries have higher wage premia (Dickens and Katz, 1987; Hasan and Chen, 2003). Empirical studies on inter-industry wage structures in the U.S. and OECD countries have found that these are relatively stable over time (see for example Krueger and Summers (1988), Zanchi (1995)). For instance, Krueger and Summers (1988) find a correlation of 0.91 between 1974 and 1984 wage premia in the US. The evidence for developing countries is mixed. While Hasan and Chen (2003) find fairly high correlations of between 0.81 and 0.84 in wage premia in the Philippines between 1988, 1994 and 1997, Goldberg and Pavcnik (2003) find that the Colombian wage premia have a low correlation over time ranging from 0.14 to 0.94 from 1984 to 1998.

The Indian wage structure is not as highly correlated over time as the US though it is not as unstable as the Colombian. There is significant variation in the magnitude of the inter-industry wage premia over time[39] but the Indian inter-industry wage structure was relatively stable during the first decade - Spearman's rank correlation coefficient is 0.9094 between 1983 and 1993. Though still fairly high the correlation of industry rankings in terms of relative wages during the 1990s was lower - Spearman's rank correlation coefficient is 0.8576 between 1993 and 1999. The rank correlation over the entire two-decade period, however, was much lower at 0.7614.[40] A Kruskal-Wallis equality of populations rank test decisively rejects the null hypothesis that the wage premia estimated in the three years are from the same population.[41] This suggests that the wage structure though initially stable became less so during the 1990s possibly due to the industrial and trade policy reforms during this period (Figure A1 in the Appendix reveals the increasing dispersion of wage premia between 1983 and 1999 from the 45-degree line). This is reflected also in the increase in the inter-industry dispersion of wages during the 1990s. The overall variability in sectoral wages is summarised using the employment-weighted standard deviation of the inter-industry wage premia adjusted for sampling variance. Again, this did not change very much during the first decade and remained around 13% between 1983 and 1993 but increased substantially to 18% in 1999.[42] This increasing dispersion coupled with the previous finding of steeper age-earnings profiles indicate a rise in the importance of unobservable skill. The extent to which the level of and the change in inter-industry wage dispersion can be attributed to the trade liberalisation of the 1990s is assessed in the following section.

6.4. The impact of trade liberalisation on relative wages

This section assesses the structure of protection as captured by the average tariff level is examined and the extent to which the trade liberalisation during the 1990s contributed to the dispersion in wages across manufacturing industries.

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[39] A test for whether wage premia estimated in one year were significantly different from those in another year was set up as follows: χ_k^2 statistic = $[\hat{\delta}_{t1}^* - \hat{\delta}_{t2}^*]'[\mathrm{var}(\hat{\delta}_{t1}^*) + \mathrm{var}(\hat{\delta}_{t2}^*)]^{-1}[\hat{\delta}_{t1}^* - \hat{\delta}_{t2}^*]$ where $\hat{\delta}_t^*$ is the (1xk) vector of the estimated wage premia in year t (t=1,2) and $var(\hat{\delta}_t^*)$ is the (kxk) variance-covariance matrix of these premia. The χ_k^2 statistics (54 degrees of freedom) are 1265.00 for 1983 and 1993, 1975.65 for 1993 and 1999, and 1008.85 for 1983 and 1999.

[40] This is true for the subset of manufacturing industries as well - the Spearman rank correlation coefficient was 0.8772, 0.8087 and 0.7614 between these pairs of years.

[41] The χ^2 statistics (2 degrees of freedom) are 6.201 for all industries and 16.90 for manufacturing industries.

[42] Similarly for manufacturing industries the adjusted employment-weighted standard deviation was about 10-11% in the first two years and rose to 14% in 1999.

6.4.1. The structure of tariff protection

Though tariff data is available for 40 agricultural, mining and manufacturing industries the subsequent analysis is undertaken for manufacturing industries only. Tariffs do not adequately capture protection in agricultural and mining industries which are subject to numerous quantitative restrictions. In the pre-reform period there were numerous controls on agricultural trade in the form of minimum export prices and quotas on agricultural exports while all agricultural imports were regulated through state agencies. During the 1990s the imports of some commodities, such as sugar and cotton, were decanalised while some edible oils were allowed on the Open General License. However imports of some commodities, such as tea, coffee and spices, were still virtually banned while others, such as natural rubber, were importable only under a license even by the end of decade (Gulati, 2000). Even as late as 1997-98, 84% of value added in agriculture was subject to a license as compared with 30% in manufacturing industries (mainly consumer goods) (Cadot et al., 2003). This suggests that using tariffs to measure protection for agriculture would not be appropriate.

Table 1 above outlined the fall in tariff and non-tariff barriers in India during the 1990s. The tariff data for manufacturing industries reveal a fall in the average tariff level from 106% to 34% and in the dispersion of tariffs across industries (the standard deviation of tariffs fell from 26% to 17%) between 1983 and 1999. This was accompanied by a fall in non-tariff barriers – the change in tariffs and NTBs have a positive correlation coefficient of 0.1946 for the two years (1993 and 1999) where there is comparable data.[43]

The structure of protection across industries has changed somewhat over time – a Kruskal-Wallis test of whether the tariff in the three years are from the same population is decisively rejected by the data.[44] The Spearman rank correlation coefficient for tariff rates for manufacturing industries is 0.5518 between 1983 and 1993, 0.6780 between 1993 and 1999

[43] This variable has not been included in the subsequent regression analysis as comparable data on non-tariff barriers for all three years is not available. Besides the common problems of quantifying and measuring NTBs there is the additional problem that only the year 1999 is common to the dataset used in this paper. The earliest available coverage ratio is for 1988-89. Given the reforms in the import licensing procedure for capital goods from 1985 onwards this figure is likely to be much lower than the actual level of NTBs in 1983. Similarly, the coverage ratio for 1995-96 (the next available year) is also likely to underestimate the level in 1993 given the reforms in the imports of agricultural commodities from 1994 onwards (Pandey, 1998).

[44] The χ_2^2 statistics are 73.84 for all tradable industries and 67.19 for manufacturing industries.

and 0.4649 between 1983 and 1999.[45] This is much lower than that intertemporal correlation in the U.S. tariffs of 0.98 between the post-Kennedy and post-Tokyo GATT rounds (1972 and 1988 respectively). Whereas Hasan and Chen (2003) find a high correlation of 0.91 between tariffs in the Philippines between 1988 and 1998, Goldberg and Pavcnik (2003) find that Colombian tariffs are only loosely correlated over time (correlations range from 0.46 to 0.94 between various year pairs).

Table A3 in the Appendix summarises the characteristics of manufacturing industries by quantiles of the tariff distribution – low (bottom 33%), medium (middle 33-66%) and high (top 33%) – in 1983. The manufacturing industries in the high tariff bracket in 1983 included the metal, plastic, rubber and chemical products; miscellaneous and woollen and silk textiles; and beverages industries. These industries have the lowest wage premia in 1983 with the lowest capital-labour ratios, skill intensity, the smallest establishment size and the lowest union density. These industries were also characterised by the lowest share of casual workers, in value added produced by registered manufacturing units, and in imports and exports. In 1983 industries in the high-tariff had the lowest share of female workers but by the 1990s the situation was exactly the opposite. All manufacturing industries were subject to licensing in 1983 but by 1999 most of the industries still requiring licenses were in the highest tariff bracket. On the other hand manufacturing industries in the highest tariff quantile had very few areas reserved for public production.

Trade liberalisation in India progressed to some extent according to end-use classification of goods. Nouroz (2001) argues that the objective of Indian trade policy was to provide protection to domestic industry and "to ensure the right kind of import structure so as to conserve scarce foreign exchange" (pp.4) and to promote industrial development. As a result, within manufacturing, tariffs were the highest for consumer goods throughout the pre- and post-reform period and lowest for capital and intermediate goods. Using the official end-use classification used for constructing the index of industrial production the industries in this dataset were divided into four groups: basic, intermediate, capital and consumer goods (EPWRF, 2003).[46] The tariff rates for consumer goods were the highest in all three years –

[45] The corresponding rank correlations for all tradable industries are 0.5724, 0.6727 and 0.6612.
[46] Basic goods included petroleum products, basic metal and heavy chemicals industries; intermediate goods included jute textiles, non-metallic mineral products, miscellaneous chemicals, paint etc, wood and paper industries, capital goods included the four machinery industries, metal products and transport equipment; and

111%, 85% and 40% - compared to capital goods – 100%, 66% and 33%. Non-tariff barriers on capital and intermediate goods were removed by 1992 and these could be freely imported on the open general license (OGL) list (Kalirajan, 2001)[47] whereas NTBs on consumer goods were reduced only in the late 1990s.

Industries in the high tariff bracket in 1999 included beverages; tobacco; soap and cosmetic products; drugs and pharmaceuticals; paints and varnishes; textiles; mineral products; other transport equipment and rubber products. Thus the most protected (in terms of tariffs) manufacturing industries in 1983 (and in the 1990s) were low-wage, relatively unskilled labour-intensive with low trade flows and produced predominantly consumer goods.

The trade reforms of the 1990s aimed to reduce both the average tariff as well as the dispersion in rates implying that the industries that were the most protected initially would have experienced the greatest declines in protection. This is borne out by the data - the industries that experienced the greatest reductions in tariffs were those with the highest initial tariff rates and were characterised by low average wages relative to other industries, high shares of unskilled labour and low shares of women and casual workers. This effect of trade liberalisation on the industry employment and wage structure is explored econometrically in the following sections.

6.4.2. Trade liberalisation and employment

At a time when tariffs fell across the board some labour reallocation from industries with higher relative tariff reductions to other industries is expected (according to the HOS trade model). This does not seem to have happened in India. There is some movement in the employment shares of regular workers but not as much as might be expected in an economy undergoing fairly drastic trade reforms. Inter-industry employment shares changed by less than one percentage point between 1983 and 1999 in all but nine industries (of which seven changed by about one to three percentage points). There is no pattern in these nine industries regarding changes in relative employment and changes in relative tariffs. A regression of the change in employment shares and change in tariff (with or without a time dummy) indicates

consumer goods included all the rest. In case of overlap the classification suggested by Nambiar et al.(1999) was followed.

[47] Some categories of capital and intermediate goods had already been placed in the OGL list in the late 1980s but these amounted to barely 30% of all imports (Panagariya, 2000).

that there is no significant relationship between the two. This provides some support for the notion that regular workers were relatively immobile in response to trade shocks.

Approaching the question from the opposite side, a discrete version of a test of the stable factor demand hypothesis, i.e. that changes in relative wages are generated solely by relative supply changes arising from reallocation of labour, was conducted between the three pairs of years. Essentially if the inequality $(W_t - W_\tau)'(X_t - X_\tau) \leq 0$ is satisfied (i.e. the inner product of the changes in wages $(W_t - W_\tau)$ and changes in factor supply $(X_t - X_\tau)$ between time period t and τ is non-positive) then supply shifts can potentially be the sole driving force for the changes in wages. "When this inequality is not satisfied, no story relying entirely on supply shifts is consistent with the data"(Katz and Murphy, 1992, pp. 48). Applying this approach to the analysis in this paper reveals that the inner product of changes in relative wages and relative labour supplies of workers for each of the three pairs of years is positive.[48] This indicates that relative demand shifts, either due to trade reforms, changes in the structure of product demand (besides through the effects of trade), or sectoral differences in factor-neutral productivity changes, have played some role in the changes in the industry wage premia. The extent of this effect that is due to trade liberalisation is explored below.

6.4.3. Trade liberalisation and the industry wage structure

In order to examine econometrically the impact of the tariff reductions during the 1990s on the industry wage structure a pooled weighted least squares (WLS) model is estimated. In the first specification wage premia are regressed on average industry tariff (expressed in fractional points). Dummy variables for the years 1993 and 1999 are included to control for time-specific macroeconomic shocks.

A bivariate plot of wage premia and tariffs for the three years - 1983, 1993 and 1999 – (see figure A2 in the Appendix) indicates a positive relationship between these two variables though there is some noise in the data. It is also clear from the figure that the two observations for industry 12 (manufacture of beverages) in the first two years have exceptionally high tariff

[48] Since trade theory focuses on relative wages as a function of relative supply and demand the inner product of relative wages (i.e. inter-industry wage premia) and relative labour supply (i.e. employment shares) rather than of average wages and employment is computed. The inner product for 1983 and 1993 is 0.0131, between 1993 and 1999 is 0.0018, and between 1983 and 1999 is 0.0027.

rates.[49] The observations for this industry in all three years were identified as influential observations in the regression analysis with high leverage values[50] and in some cases high residual values as well. The exclusion of these observations had an impact on the coefficient on tariffs. Though some other observations were identified as possible outliers (e.g. manufacture of tobacco in 1993) omitting these observations from the dataset did not influence the estimated coefficients. The specifications estimated below do not include the beverages industry.

To the extent that political economy considerations and the ranking of tariffs on the basis of unobservable worker characteristics (e.g. protection is higher in industries with lower wages or with less skilled workers) are important there is a danger of spurious correlation between tariffs and relative wages. In order to mitigate this problem of potential simultaneity bias alternative specifications are estimated with controls for determinants of inter-industry wage premia other than tariffs.[51]

The second specification includes variables indicating the average level of observable human capital as suggested by Dickens and Katz (1987). Though the effect of individual skill (proxied by education) is captured in the wage regression models it is possible that the average level of skill in an industry also influences the average wage in that industry.[52] They also argue that unobserved labour quality or discrimination might crowd female workers into certain jobs that reduce their wages and that this crowding also reduces the wages of any other workers in that job. If this is so then the share of female workers in the industry's workforce would influence the average wage in that industry. In this paper the share of skilled, female and casual workers in an industry are included in the second specification. In addition, Aghion et al.(2003) find that the impact of trade liberalisation on productivity growth and

[49] In all three years the industry tariff rate is extremely high – 193%, 224% and 123% compared to an average rate of 96%, 70% and 31% in 1983, 1993 and 1999 respectively for the other sectors – mainly due to the high tariff on goods with alcohol content.

[50] Velleman and Welsch (1981) recommend using $3\,p/n$ (where p is the number of variables and n is the number of observations) as the calibration point in small samples.

[51] The frequency rate of fatal injuries reported by establishments in the organised sector was also included in the regression models. The hypothesis is that wage premia reflect in part compensatory differentials paid to workers to compensate them for non-pecuniary aspects of their work. This variable was ultimately dropped as it did not have a significant effect on wage premia in any of the specifications. Other empirical studies also do not give much support to the compensatory differentials approach (Krueger and Summers, 1987).

[52] This might be because more able workers might be attracted to industries paying higher wages or because workers in industries where other workers are well educated or more experienced may also be better workers.

profits differed across states depending on their labour market institutions.[53] In order to capture this regional dimension at the industry level the share of employment in total industry employment in pro- and anti-worker state groups (following Besley and Burgess (2002) are included in the second specification. The hypothesis is that workers in industries concentrated in pro-worker states would have greater bargaining power that possibly enables them to extract higher wages (or employment guarantees).

Jean and Nicoletti (2002) point out that market structure also plays an important role in determining relative wages. Efficiency wage models predict that firms operating in less competitive markets and facing a relatively inelastic product demand curve would have greater ability to pay workers higher than competitive wages. As a result studies of determinants of wage premia usually include variables that capture the degree of competition either through industrial concentration, barriers to entry or degree of import penetration. The first of these variables is not included in the analysis in this paper for three reasons: one, it is difficult to obtain comparable data on industry concentration ratios for all three years;[54] two, the variables used to capture industrial regulation are likely to be highly correlated with any measure of industrial concentration (see below for a more detailed explanation); and three, industrial regulation measures are preferable as they provide a direct link to policy unlike measures of concentration or product market rents (Jean and Nicoletti, 2002). Capturing the effect of barriers to entry are particularly important in the Indian context as the trade reforms in the 1990s were preceded and accompanied by domestic industrial deregulation. The third specification includes two dummy variables to capture industrial policy reform with respect to licensing and reservation for the public sector.

The fourth specification further includes other determinants of relative wages such as unionisation,[55] establishment size and capital intensity. The union threat model proposed by

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[53] Both central and state governments are empowered to introduce legislation on matters related to trade unions as well as industrial and labour disputes implying that firms located in different states might face different regulatory climates (Besley and Burgess, 2002).

[54] The data on the Herfindahl index provided by Dr. Uma Kambhampati is available for a sample of medium and large non-government, non-financial, public limited companies collected by the Reserve Bank of India (Kambhampati and Kattuman, 2003). These data could not be used in this paper as the mapping of industries in that dataset to that used in this paper was necessarily ad hoc with no data on three industries; the most recent year in the concentration dataset was 1997; and there are some concerns that these data apply to the specific sample only and not the entire population. Dickens and Katz (1987) point out that using a concentration measure at an aggregation level not closely related to the relevant product market would tend to bias estimates.

[55] Studies on the determinants of inter-industry premia often include factors like union membership or coverage and establishment size in the first-stage wage regression models and estimate inter-industry wage premia after

Dickens (1986) argues that firms may pay high wages to avoid the threat of collective action. As a result industries where conditions are favourable to unionisation (e.g. larger plant sizes), workers are predisposed to form unions and the firms have the ability to pay or have higher profits to share would be more likely to have higher wages (Jean and Nicoletti, 2002). Most empirical studies have found a positive effect of industry union density on the wages of both unionised and non-unionised workers (though this is sensitive to the data and methodology used) (Dickens and Katz, 1987). The average establishment size has usually been found to be positively related to the average industry wage though this effect is more important in determining intra- rather than inter-industry wage premia (Brown and Medoff, 1989).

Efficiency wage models indicate that capital intensity is also likely to be positively related to worker bargaining power and wages. Essentially, for industries with a relatively high capital-labour ratio the cost of raising wages is lower (Jean and Nicoletti, 2002). The problem here is the possibility that industry-specific capital could respond endogenously to changes in labour costs, i.e. it is not possible to determine whether capital intensive industries need to pay higher wages or whether a high wage premium arising due to other factors results in the substitution of capital for labour (Dickens and Katz, 1987). The capital-labour ratio is also likely to be correlated with the regional labour market variables – industries concentrated in pro-worker states may substitute labour for capital. As a result the capital-output ratio instead of the capital-labour was included in the fourth specification more to check the robustness of the results than to analyse the role of capital intensity in determining relative wages. Though the potential simultaneity bias still exists it is likely to be less of a problem.

The results of these four specifications are reported in Table 4 below. The first four regression models are estimated using Weighted Least Squares (WLS) with the variance in the estimated wage premia as the weights. The last specification is also estimated using Ordinary Least Squares (A4') as a robustness test.

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controlling for these factors (see for example Brown and Medoff (1989)). This data are not available for all 3 years at the individual level in the Indian survey data (and there are many missing values in the years for which the data are available) so average industry data is used in these second-stage regression models.

Table 4: Determinants of wage premia for manufacturing industries

Dependent variable: estimated inter-industry wage premia [a]

	Weighted Least Squares (WLS)				OLS	
	A1	A2	A3	A4	A4'	Mean
Average tariff [a]	0.1309	0.2016***	0.1938***	0.2456***	0.1360**	0.7066
	(0.0793)	(0.0624)	(0.0629)	(0.0580)	(0.0666)	
Time dummy for 1993	-0.0037	-0.0202	-0.0001	0.0318	-0.0327	0.33333
	(0.0356)	(0.0274)	(0.0332)	(0.0310)	(0.0342)	
Time dummy for 1999	-0.0296	-0.0217	-0.0021	0.0255	-0.0676	0.33333
	(0.0621)	(0.0483)	(0.0515)	(0.0469)	(0.0529)	
Share of female workers in industry		-0.1503	-0.0792	-0.0852	0.0115	0.1566
		(0.0909)	(0.0980)	(0.0887)	(0.0868)	
Share of casual workers in industry		0.1148	0.1250	0.1395	-0.0108	0.1549
		(0.1214)	(0.1201)	(0.1081)	(0.1140)	
Share of skilled workers in industry		0.3448***	0.3451***	0.4424***	0.4933***	0.3361
		(0.0712)	(0.0710)	(0.0681)	(0.0683)	
Pro-worker states [b]		0.2078**	0.1800*	-0.0303	-0.1189	0.3818
		(0.0943)	(0.0944)	(0.1010)	(0.1006)	
Anti-worker states [b]		0.0562	-0.0531	-0.0310	-0.1122	0.3832
		(0.1325)	(0.1424)	(0.1341)	(0.1273)	
Industrial license dummy			0.0153	0.0155	0.0080	0.4839
			(0.0286)	(0.0259)	(0.0285)	
Public sector dummy			0.0538*	0.0557**	0.0555*	0.1290
			(0.0281)	(0.0262)	(0.0300)	
Union density				0.1110**	0.0804	0.1883
				(0.0442)	(0.0589)	
Average establishment size				0.0003***	0.0003***	91.5406
				(0.0001)	(0.0001)	
Capital-output ratio				0.0314	0.0828**	0.3593
				(0.0292)	(0.0395)	
Constant	-0.1163	-0.3534***	-0.3365***	-0.4217***	-0.2436**	
	(0.0835)	(0.1130)	(0.1129)	(0.1070)	(0.1113)	
Observations	93	93	93	93	93	
Adjusted R-squared [c]	0.1099	0.4776	0.5021	0.5994	0.5888	

Note: The beverages industry has been excluded from the dataset. a\ Wage premia and tariffs are expressed as proportional points. Mean wage premia is -0.0180. b\ Concentration of employment in neutral states, i.e. states with neither pro- or anti-worker amendment activity, is the omitted category (mean 0.2350). c\ For the WLS specifications: as the reported R^2 does not provide a useful measure of goodness of fit of the original (unweighted) model the adjusted R^2 listed in the table is computed as the squared correlation between the actual and predicted values of the dependent value (where the predicted value is computed using the unweighted data and efficient parameter estimates from the WLS regression model) (Rubinfeld and Pindyck, 2001).

Tariff effect

The tariff-wage effect is consistently positive and significant at the 1% level or better across all except the first specification (where it is significant at the 10.2% level), indicating that industries with relatively higher protection would tend to have relatively higher wages. The magnitude of this effect is of a similar magnitude across all specifications (except the first), i.e. a 10% fall in the tariff in an industry would lead to a 2-2.5% fall in the wage premium of that industry. For example, moving a worker from an industry with the average level of protection in 1983 (106% tariff rate) to an industry with zero tariff the estimates in the last specification (A4) implies a substantial fall of about 26% (0.2456* 1.06 * 100) in the wage premium. The OLS estimate of the tariff effect reported in specification A4' is also positive though smaller – a 10% fall in the industry tariff will reduce the wage by 1.4%.

This positive relationship between tariff protection and relative wages is consistent with the Ricardo-Viner trade model under the assumption that labour is immobile across industries. This models postulates that protection of an industry reduces competition from imports and raises the demand for labour which in turn raises the wages in that industry as labour reallocation across industries is constrained and does not eliminate the differential. The existence of labour market rigidities seems relevant in the Indian context where employment security is of paramount importance for workers (Ghose, 1995). Data from the 1999 NSS survey reveals that there is practically no mobility of workers across industries - only about 1% of all workers and 1.2% of male regular workers had changed industries in the two years preceding the survey.[56] Workers with regular wage or salaried jobs presumably build up firm- and industry-specific capital or acquire seniority status that makes them less likely to move across industries in response to shocks to labour demand.

This positive tariff effect is also consistent with the notion that trade reduces distortions in an imperfectly competitive market thereby eroding rents and leading to a fall in relative wages. Additionally, if there are improvements in productivity following trade liberalisation that occur differentially across industries and these improvements are passed on to labour then tariff reductions would result in higher relative wages.

In contrast, Gaston and Trefler (1994) find a significant negative effect of tariffs on relative wages in U.S. manufacturing industries in 1983 even after treating protection as endogenous.[57] Possible explanations for this negative correlation include the possibility that unions take advantage of protection by offering wage concessions in exchange for employment guarantees. Another possible reason is that efficient reallocation of labour from

[56] This question was not asked in the previous two surveys. However a crude approximation for migrant workers in 1983 revealed that about 3% of these migrants had changed their industry over the past two years preceding the survey.
[57] Their regression model also includes various other trade-related variables such as non-tariff barriers (NTBs), exports, imports, import growth over the previous three years and intra-industry trade. Though NTBs are found to be associated with higher wages this effect is not statistically significant in most specifications.

import-competing sectors has been discouraged by long-term protection and the negative tariff-wage effect simply reflects this inefficiency.

Goldberg and Pavcnik (2003), on the other hand, estimate a positive correlation between tariff protection and industry wage premia after controlling for unobserved sectoral heterogeneity and political economy factors through industry fixed effects.[58] They regress changes in wages on changes in tariff and estimate that a 10% reduction in tariff would decrease relative wages by 1.4%. This is similar to the estimated tariff effect in the last three specifications (see Table 4 above). This positive tariff-wage relationship was found to be robust to the inclusion of trade flow and other industry-specific variables and also to instrumenting for trade policy changes. Hasan and Chen (2003) estimate a negative tariff-wage effect (with controls for lagged trade flows and exchange rate) for the Philippines that becomes insignificant once industry fixed effects are included. Estimation of inter-industry wage premia across 12 OECD countries including the U.S. using panel data for 1996 reveals a strong positive impact of tariff on relative wages in manufacturing industries (Jean and Nicoletti, 2002). This specification includes other trade variables such as NTBs, import penetration and export intensity, industry-specific variables such as union density and establishment size as well as industry and country dummies. On the whole it seems that the direction of the effect of trade policy on the wage outcome as captured by average tariff rates is largely positive for manufacturing industries.

The time dummies are not significant at a conventional level indicating that time-variant and industry-invariant factors do not play a significant role in determining wage premia.

Unobserved labour quality

The shares of casual, skilled and female workers in total wage employment are included in the equations to control for the effect of unobserved labour quality or discrimination on the average industry wage. The share of casual workers does not have a significant impact on wages possibly due to the low participation of these workers in manufacturing. The share of skilled workers in an industry has a significantly positive effect on the average industry wage - increasing the skill intensity in an industry by one percentage point would raise the average wage by approximately 0.4 of a percentage point. Dickens and Katz (1987) also find that the

[58] They estimate a negative tariff-wage effect in manufacturing industries when there are no controls for industry fixed effects.

effect of the average observable human capital as measured by the years of education is positive and significant in each of their numerous specifications for the US. This suggests that productivity would be higher in an industry with a greater proportion of skilled workers, over and above the marginal product of labour (as captured by the effect of individual skill levels on the individual wage).

The effect of high female employment share exerts a downward pressure on industry wages though this effect is not significant. Dickens and Katz (1987) also estimate a negative relationship between feminisation and relative wages in the U.S. The experience of several Asian countries has been that deregulation and liberalisation has generated increased demand for female labour, particularly in light manufacturing industries and export-oriented industries, as women are often willing to accept lower wages, temporary contracts and inferior working conditions and have lower unionisation rates (Kanji and Menon-Sen, 2001).[59] In India the share of female employment varies considerably across industries and women are predominantly employed in light manufacturing industries. During the 1990s the share of female employment rose in these industries, especially in the manufacturing of beverages, tobacco, textiles (except jute textiles), wood products, as well as in chemicals industries.

State labour market regulations

Differences in state labour market regulations influence the inter-industry wage premia – increasing the share of industry employment in pro-worker states (relative to the employment share in neutral states) by one percentage point would raise the relative industry wage by about 0.2 of a percentage point. Aghion et al. (2003) find that pro-worker legislations discouraged productivity growth, profitability, employment and output and encouraged the substitution of capital for labour in the organised manufacturing sector between 1980 and 1997. While low profitability would tend to lower wages, low employment would tend to raise wages and low labour share in total costs (due to the high capital intensity) would make it easier for firms to pay higher than competitive wages. Unsurprisingly, these state effects become insignificant once controls for union density, establishment size and capital-output ratio are included (specification A4).

[59] A case study of the knitwear industry confirms this pattern in India (Neetha, 2002).

Industrial policy effect

The industrial license and public dummy variables capture the effects of industrial policy on relative wages through their effect on the production function as well as the market structure. The license dummy has a positive but insignificant effect on the average wage premia, possibly due to the lack of inter-industry variation in 1983 as all manufacturing industries had to obtain licenses for entry and expansion.[60] The uncertainty, time and lobbying associated with obtaining licenses restricted competition and generated rents for producers with licenses. There is some anecdotal evidence of deliberate pre-emption of capacity and that the licensing policy favoured large business houses thereby limiting competition (Sengupta, 1992).

The reservation of an industry for production by the public sector raises the industry wage premia by about 0.1 of a percentage point.[61] In 1983, 8 of the 31 manufacturing industries (excluding beverages) had items that were reserved for production by the public sector only; these included the metal (iron and steel), all machinery, transport equipment, miscellaneous chemicals and miscellaneous manufacturing industries. By 1993 only two of these – miscellaneous chemicals (arms and ammunitions) and transport equipment (railway) – were reserved and these remained so in 1999 as well. Though comparable data on concentration ratios for all three years are difficult to obtain, firm-level data for 1987-88 reveals that the four-firm concentration ratio for industries reserved for the public sector was as high as 77.9 (Desai, 1992). The public reservation dummy are likely to be highly correlated with any measure of industrial concentration as the one of the reasons for reserving certain sectors was that the private sector would not be able to raise the investment required. Thus the positive public premia reflects the preferential treatment given by the government to these sectors that were identified as critical to the development of the country.

Other industry-specific variables

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[60] By 1993 about 11 of the 31 industries in this dataset required licenses (such as beverages, wood and wood products, paper and paper products, leather and leather products, drugs and medicines, miscellaneous chemicals, basic metal industry, etc.). These industries are very diffuse in terms of worker and industry characteristics – the capital-output ratio varies from 0.10 in the tobacco industry to 0.87 in the basic metal industry while the share of skilled workers varies from 0.85 in the pharmaceuticals industry to 0.06 in the tobacco industry.

[61] It should be noted that this variable captures the effect of product market regulation. It is not an estimate of a possible public sector wage premia accruing to workers employed in public sector enterprises. In fact workers in public sector enterprises are paid less than in private sector enterprises – though this comparison is made without controlling for individual characteristics. A dummy variable for public sector affiliation could not be included in the wage regression models as only the last two employment surveys report whether or not a worker is employed in the public sector and there are many missing values.

The coefficients on the additional industry-specific factors - union density, average establishment size and capital-output ratio - estimated in the fourth specification (A4) all have the expected positive effect on the wage premia.

Union density has a significant positive effect on wages – a one percent increase in this variable raises the average industry wage by about 0.1% Though of the expected sign this is quite low in magnitude compared to union wage gap estimates for the US (about 15%), UK (10%), Canada (10-15%), Australia (7-17% with most estimates at the lower end of the range) (Blanchflower, 1997). These estimates, however, are all based on the wage regression models from microeconomic data. It is possible that the average industry union density would have a lower effect on wages for both unionised and non-unionised workers. There are also concerns about the union membership data as mentioned earlier that possibly understate the true extent of unionisation.[62] There is some evidence that employment security is of great importance to Indian workers and the majority of labour regulations as well as industrial disputes focus on this issue (Ghose, 1995). It is possible that unions seek employment guarantees even at the cost of wage increases. Fallon and Lucas (1993) find that (until the 1980s) the fall in labour demand following the enactment of labour regulations was lower in industries with high union density. In addition, Nagaraj (1994) claims that "in a period [1980s] of declining bargaining power of organised workers and structural changes in employment within registered sector towards smaller sized establishments, unionised labour is unlikely to have secured a disproportionate increase in the wage rate" (pp. 180).

The average establishment size in India is quite small – about 92 employees per factory in the organised sector. This could be a consequence of excess entry due to protection or because firms prefer small sizes due to the labour market regulations.[63] Table 4 indicates that increasing the average establishment size raises the industry wage premia by about 3% (=coefficient times mean establishment size). This is similar to the estimates of the establishment size-wage effect from individual-level data from the U.S. surveyed in Brown and Medoff (1989) – these range from 0.8% to 3.8%. They find that unobservable indicators of labour quality are a possible explanation for this positive effect.

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[62] Only about 18% of registered unions submitted returns that are reported in the trade union data in 1983. This fell further to 12.2% by 1993 (Bhattacherjee, 1999).

[63] The rules for lay-off, retrenchment and closure under the central Industrial Disputes Act (amended in 1982) apply to establishments with 100 or more workers but some state governments, e.g. West Bengal, have amended this to apply to 50 or more workers (Besley and Burgess, 2002).

The capital-output ratio has a significantly positive effect on wages as expected of a measure of labour productivity. In addition, in the presence of imperfect competition a low labour share implies a lower cost to raising wages above competitive levels. However, as mentioned earlier concerns that this variable is potentially endogenous imply that it serves more as a conditioning variable for the tariff-wage effect rather than to analyse the role of capital intensity in determining relative wages.

While the first specification explains only about 11% of the variation in the industry wage premia the preferred specification (A4 - WLS) explains almost 60% of this variation. As the dependent variable, wage premia, is estimated from the wage equation the WLS models that give lower weight to observations with large variance are preferred. As a further sensitivity test trade flows are included in the specifications. To the extent that informal trade barriers, non-tariff barriers, differential transport and communications costs, and other industry-specific barriers affect trade flows these variables will capture the overall effect of all trade-related channels, other than trade policy that is captured by tariffs, on wages. Unlike Gaston and Trefler (1994) who estimate a significant negative and positive effect of imports and exports respectively on industry wages, Goldberg and Pavcnik (2003) find no significant relationship between lagged import and export flows and industry wages.

Measures of trade (industry-level real import and export flows in constant 1983 prices, import and export shares) are included (separately) in all four specifications described above. The effect of the import variables (real import flow or share) on the industry wage is positive while that of the export variables is negative (though not always significant). These effects are consistent with the notion of comparative advantage. Countries tend to export goods that use intensively the relatively abundant factor; in India this is unskilled low-wage labour. The pattern of trade indicates that industries with high import flows or shares were in general intermediate and capital goods industries, such as the base metal industry, electrical and non-electrical machinery, petroleum products, chemicals etc.[64] Many of these industries were identified as critical sectors for the Indian economy and had some sub-sectors reserved for public sector production. It is unlikely that India would have a comparative advantage in the production of these goods. However, as these trade flow variables are potentially endogenous

[64] An exception is the edible oils industry which has high imports but these were canalised until 1995, i.e. all imports were through state agencies.

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 to the inclusion of these trade variables – the tariff coefficient varies in a tight range of –0.03 and +0.02 around the values reported in Table 4.

7. Conclusion

This paper represents one of the first attempts at analysing econometrically the link between trade protection and inter-industry wage premia in India. This analysis combines detailed tariff data with microeconomic survey data for three years - 1983, 1993 and 1999 - that span the period of rapid trade liberalisation in the 1990s. While the first year reflects the structure of the Indian labour markets prior to the reforms the latter two provide the basis for analysing the impact of the reforms on these structures. Inter-industry wage premia are estimated using information on worker characteristics after controlling for potential selection bias. This paper finds that there is substantial dispersion of wages across industries and that there has been some change in the industry wage structure over time.

The major finding of this paper is that the impact of trade liberalisation on the inter-industry wage premia for regular workers is substantial and that more protected industries tend to have higher relative wages. Conversely, industries that undergo tariff reductions have lower wages relative to other industries. This positive tariff-wage effect is evident whether or not industry fixed effects, such as worker composition, regional concentration, industrial policy, union density, average establishment size and capital intensity, are included. This positive effect could reflect the erosion of rents that are received (and reflected in the wages earned) by unionised workers in imperfectly competitive markets following trade liberalisation. It is also consistent with the short-run specific factors and the medium-run Ricardo-Viner models of trade that predict a positive relationship between tariffs and inter-industry wage premia.

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APPENDIX

Table A1: Summary statistics : Wage regression models

	1983		1993		1999	
	Mean	**S.D.**	**Mean**	**S.D.**	**Mean**	**S.D.**
Wage data (Rs.):						
Natural log of real hourly wages	1.2679	0.5176	1.5173	0.5867	1.6919	0.6539
Individual characteristics:						
Age	34.8790	11.0472	36.7696	10.8995	37.1110	11.2464
Age splines:						
Age: 15-25 years	24.1235	2.1711	24.3587	1.8793	24.3413	1.8730
Age: 25-35 years	6.2395	4.2634	6.9669	4.0333	6.9569	4.0819
Age: 35-45 years	3.2483	4.2377	3.9034	4.3920	4.0739	4.4414
Age: 45-55 years	1.1109	2.7388	1.4049	2.9901	1.5646	3.1579
Age: 55-65 years	0.1570	0.9775	0.1357	0.8399	0.1742	0.9289
Married	0.7760	0.4170	0.8030	0.3978	0.7889	0.4081
Education:						
Literate ‡	0.0987	0.2982	0.0823	0.2748	0.0715	0.2577
Completed primary school	0.1472	0.3543	0.1027	0.3035	0.0969	0.2958
Completed middle school	0.1745	0.3795	0.1612	0.3677	0.1694	0.3752
Completed secondary school	0.2560	0.4364	0.3106	0.4627	0.3300	0.4702
Completed graduate school	0.1403	0.3473	0.2342	0.4235	0.2408	0.4276
Social exclusion:						
Member of scheduled caste or tribe	0.1770	0.3816	0.1555	0.3624	0.1793	0.3836
Muslim	0.0937	0.2914	0.0888	0.2845	0.1024	0.3032
Other ‡	0.7293	0.4443	0.7556	0.4297	0.7183	0.4498
Seasonality:						
Household interviewed in:						
1st quarter ‡	0.2648	0.4412	0.2543	0.4355	0.2528	0.4346
2nd quarter	0.2519	0.4341	0.2486	0.4322	0.2519	0.4341
3rd quarter	0.2384	0.4261	0.2459	0.4306	0.2464	0.4309
4th quarter	0.2449	0.4300	0.2512	0.4337	0.2489	0.4324
Rural * 1st quarter ‡	0.0820	0.2744	0.0722	0.2588	0.0758	0.2647
Rural * 2nd quarter	0.0775	0.2674	0.0718	0.2581	0.0707	0.2563
Rural * 3rd quarter	0.0725	0.2593	0.0714	0.2574	0.0683	0.2522
Rural * 4th quarter	0.0778	0.2679	0.0696	0.2545	0.0701	0.2554
Location:						
Residence in rural areas	0.3098	0.4624	0.2850	0.4514	0.2849	0.4514
State of residence:						
Andhra Pradesh	0.0808	0.2726	0.0837	0.2770	0.0816	0.2738
Assam	0.0445	0.2063	0.0372	0.1892	0.0430	0.2030
Bihar	0.0603	0.2380	0.0518	0.2216	0.0437	0.2044
Gujarat	0.0579	0.2336	0.0614	0.2401	0.0591	0.2359
Haryana	0.0184	0.1342	0.0177	0.1320	0.0204	0.1414
Himachal Pradesh	0.0136	0.1160	0.0199	0.1395	0.0311	0.1735
Karnataka	0.0507	0.2195	0.0539	0.2258	0.0530	0.2240
Kerala	0.0369	0.1885	0.0343	0.1820	0.0379	0.1909
Madhya Pradesh	0.0775	0.2674	0.0764	0.2656	0.0651	0.2466
Maharashtra	0.1514	0.3584	0.1509	0.3580	0.1470	0.3541
Orissa	0.0318	0.1754	0.0329	0.1784	0.0299	0.1704
Punjab	0.0516	0.2212	0.0487	0.2152	0.0511	0.2201
Rajasthan	0.0351	0.1839	0.0411	0.1985	0.0491	0.2161
Tamil Nadu	0.0876	0.2828	0.0844	0.2780	0.1017	0.3023

	1983		1993		1999	
	Mean	**S.D.**	**Mean**	**S.D.**	**Mean**	**S.D.**
Tripura	0.0088	0.0934	0.0191	0.1370	0.0138	0.1166
Uttar Pradesh	0.0924	0.2896	0.0904	0.2867	0.0965	0.2953
West Bengal ‡	0.1007	0.3009	0.0963	0.2950	0.0759	0.2649
Industry affiliation:						
Agricultural and allied activities:						
Food crops	0.0774	0.2672	0.0327	0.1780	0.0332	0.1792
Cash crops	0.0032	0.0563	0.0014	0.0379	0.0014	0.0368
Plantation crops	0.0199	0.1396	0.0089	0.0938	0.0147	0.1203
Other crops	0.0009	0.0296	0.0011	0.0331	0.0011	0.0326
Animal husbandry	0.0078	0.0879	0.0041	0.0636	0.0030	0.0544
Forestry and fishing	0.0019	0.0431	0.0025	0.0500	0.0038	0.0613
Mining sector:						
Fuels	0.0157	0.1244	0.0188	0.1357	0.0152	0.1225
Minerals	0.0049	0.0696	0.0043	0.0656	0.0023	0.0476
Manufacturing sector:						
Sugar products	0.0045	0.0672	0.0047	0.0681	0.0053	0.0724
Edible oils	0.0026	0.0505	0.0025	0.0503	0.0011	0.0337
Miscellaneous food products	0.0105	0.1019	0.0132	0.1141	0.0119	0.1083
Beverages	0.0016	0.0401	0.0020	0.0448	0.0016	0.0397
Tobacco	0.0073	0.0854	0.0039	0.0627	0.0019	0.0440
Cotton textiles	0.0341	0.1816	0.0225	0.1485	0.0199	0.1395
Woollen and silk textiles	0.0065	0.0802	0.0140	0.1176	0.0067	0.0818
Jute textiles	0.0091	0.0948	0.0077	0.0872	0.0033	0.0570
Miscellaneous textile products	0.0160	0.1257	0.0127	0.1118	0.0151	0.1219
Wood products incl. furniture	0.0071	0.0839	0.0047	0.0681	0.0049	0.0702
Paper products	0.0033	0.0569	0.0039	0.0624	0.0043	0.0656
Printing and publishing	0.0086	0.0921	0.0072	0.0848	0.0324	0.1771
Leather and leather products	0.0034	0.0582	0.0039	0.0621	0.0060	0.0775
Rubber products	0.0026	0.0509	0.0034	0.0586	0.0038	0.0613
Plastic products	0.0019	0.0436	0.0029	0.0539	0.0045	0.0670
Petroleum products	0.0011	0.0331	0.0024	0.0492	0.0025	0.0502
Heavy chemicals	0.0034	0.0585	0.0049	0.0700	0.0023	0.0484
Fertilisers and pesticides	0.0018	0.0427	0.0030	0.0543	0.0034	0.0583
Paints, varnishes and lacquers	0.0017	0.0414	0.0018	0.0426	0.0022	0.0468
Drugs and medicines	0.0035	0.0594	0.0055	0.0742	0.0052	0.0717
Soaps, cosmetics and perfumes	0.0023	0.0476	0.0020	0.0448	0.0012	0.0348
Miscellaneous chemicals	0.0048	0.0693	0.0052	0.0716	0.0026	0.0506
Non-metallic mineral products	0.0116	0.1070	0.0106	0.1026	0.0100	0.0995
Basic metal industry	0.0229	0.1497	0.0206	0.1421	0.0158	0.1245
Metal products	0.0103	0.1012	0.0108	0.1035	0.0153	0.1227
Machinery for agriculture and food and textile industries	0.0023	0.0483	0.0028	0.0532	0.0023	0.0484
Other non-electrical machinery	0.0118	0.1079	0.0156	0.1240	0.0122	0.1096
Electrical industrial machinery	0.0070	0.0833	0.0050	0.0708	0.0063	0.0791
Electrical appliances and electronics	0.0078	0.0881	0.0079	0.0886	0.0091	0.0949
Transport equipment -sea, rail and motor vehicles	0.0173	0.1304	0.0145	0.1196	0.0164	0.1272
Other transport equipment	0.0038	0.0618	0.0041	0.0641	0.0034	0.0586
Misc. manufacturing industry	0.0105	0.1019	0.0177	0.1320	0.0148	0.1208
Utilities:						
Electricity	0.0192	0.1371	0.0205	0.1418	0.0188	0.1358
Gas & water supply	0.0030	0.0547	0.0034	0.0580	0.0043	0.0653

	1983		1993		1999	
	Mean	**S.D.**	**Mean**	**S.D.**	**Mean**	**S.D.**
Construction	0.0174	0.1309	0.0166	0.1278	0.0182	0.1338
Wholesale and retail trade	0.0585	0.2347	0.0694	0.2541	0.1048	0.3063
Hotels and restaurants	0.0188	0.1357	0.0171	0.1298	0.0224	0.1479
Transport, storage and communication:						
Railway transport services	0.0437	0.2044	0.0349	0.1835	0.0256	0.1580
Other transport services, Storage and warehousing	0.0578	0.2334	0.0608	0.2390	0.0747	0.2629
Communication	0.0143	0.1188	0.0121	0.1093	0.0177	0.1318
Services:						
Banking	0.0292	0.1683	0.0368	0.1882	0.0340	0.1813
Insurance	0.0029	0.0540	0.0044	0.0664	0.0042	0.0645
Education and research	0.0825	0.2751	0.1026	0.3034	0.1012	0.3016
Medical and health services	0.0218	0.1460	0.0221	0.1470	0.0237	0.1520
Other services [a]	0.0446	0.2064	0.0526	0.2233	0.0372	0.1892
Public administration	0.2117	0.4085	0.2289	0.4201	0.1908	0.3930
Selectivity bias correction term	1.1489		1.1256		1.1740	
Total number of observations	27356		26387		27295	

Notes: S.D. stands for standard deviation of the variable. In the case of dummy variables the mean refers to the percentage of observations falling in each category. Some observations (amounting to about 1-2% in the three years) had to be dropped from the sample as there were missing observations on wages, hours worked and industry affiliation. ‡ indicates omitted dummy variable. Notes: In the case of dummy variables the mean refers to the percentage of observations falling in each category. a\ Other services comprises legal, business, personal, social, sanitary and community services.

Table A2: Inter-industry wage premia, 1983-1999

	Wage equation coefficients			Wage premia		
	1983	**1993**	**1999**	**1983**	**1993**	**1999**
Agricultural and allied activities:						
Food crops	0.4520***	0.6989***	0.8341***	-0.2198***	-0.2533***	-0.2096***
	(0.0571)	(0.0911)	(0.0832)	(0.0079)	(0.0131)	(0.0154)
Cash crops	0.4239***	0.6192***	0.7168***	-0.2479***	-0.3330***	-0.3269***
	(0.0654)	(0.1036)	(0.1047)	(0.0318)	(0.0522)	(0.0653)
Plantation crops	0.4725***	0.5988***	0.5362***	-0.1993***	-0.3534***	-0.5074***
	(0.0588)	(0.0933)	(0.0838)	(0.0132)	(0.0237)	(0.0197)
Other crops	0.5960***	0.8302***	0.7639***	-0.0758	-0.1221*	-0.2798***
	(0.0926)	(0.1131)	(0.1028)	(0.0697)	(0.0699)	(0.0662)
Animal husbandry	0.3921***	0.7240***	0.7694***	-0.2797***	-0.2283***	-0.2743***
	(0.0598)	(0.0948)	(0.0900)	(0.0230)	(0.0370)	(0.0454)
Forestry and fishing	0.5567***	0.9351***	1.1135***	-0.1151***	-0.0171	0.0698*
	(0.0704)	(0.0961)	(0.0907)	(0.0413)	(0.0413)	(0.0421)
Mining sector:						
Fuels	1.0301***	1.2559***	1.4174***	0.3583***	0.3036***	0.3737***
	(0.0586)	(0.0893)	(0.0831)	(0.0161)	(0.0217)	(0.0175)
Minerals	0.7839***	0.9823***	0.9708***	0.1121***	0.0300	-0.0729
	(0.0638)	(0.0973)	(0.0959)	(0.0297)	(0.0387)	(0.0497)
Manufacture of :						
Sugar products	0.7594***	0.9757***	0.9537***	0.0876***	0.0235	-0.0900**
	(0.0671)	(0.0959)	(0.0884)	(0.0338)	(0.0359)	(0.0375)
Edible oils	0.5386***	0.7840***	0.8086***	-0.1332***	-0.1682***	-0.2351***
	(0.0678)	(0.0991)	(0.1048)	(0.0386)	(0.0422)	(0.0651)

	Wage equation coefficients			Wage premia		
	1983	1993	1999	1983	1993	1999
Miscellaneous food products	0.5815***	0.8210***	0.9223***	-0.0903***	-0.1313***	-0.1214***
	(0.0590)	(0.0920)	(0.0831)	(0.0199)	(0.0210)	(0.0215)
Beverages	0.6722***	0.9502***	0.9720***	0.0005	-0.0020	-0.0717
	(0.0691)	(0.1144)	(0.1024)	(0.0395)	(0.0656)	(0.0635)
Tobacco	0.4385***	0.5499***	0.9111***	-0.2333***	-0.4024***	-0.1326**
	(0.0611)	(0.0976)	(0.0999)	(0.0226)	(0.0411)	(0.0616)
Cotton textiles	0.7078***	0.8164***	0.8878***	0.0360***	-0.1358***	-0.1559***
	(0.0572)	(0.0906)	(0.0827)	(0.0114)	(0.0144)	(0.0197)
Woollen and silk textiles	0.6673***	0.9654***	0.9086***	-0.0044	0.0131	-0.1351***
	(0.0598)	(0.0923)	(0.0873)	(0.0227)	(0.0199)	(0.0321)
Jute textiles	0.7621***	1.0271***	0.9293***	0.0903***	0.0748***	-0.1144***
	(0.0577)	(0.0891)	(0.0868)	(0.0187)	(0.0171)	(0.0353)
Misc. textile products	0.6274***	0.8297***	0.8980***	-0.0444**	-0.1226***	-0.1456***
	(0.0574)	(0.0896)	(0.0833)	(0.0173)	(0.0208)	(0.0199)
Wood products	0.5991***	0.8695***	0.8459***	-0.0727***	-0.0827**	-0.1978***
	(0.0601)	(0.0934)	(0.0889)	(0.0222)	(0.0336)	(0.0373)
Paper products	0.7817***	1.0072***	0.9600***	0.1099***	0.0549	-0.0837**
	(0.0677)	(0.0960)	(0.0888)	(0.0383)	(0.0372)	(0.0349)
Printing and publishing	0.5540***	0.7737***	0.8304***	-0.1178***	-0.1785***	-0.2133***
	(0.0605)	(0.0946)	(0.0811)	(0.0218)	(0.0287)	(0.0158)
Leather and leather products	0.6767***	0.8867***	0.9119***	0.0049	-0.0656**	-0.1318***
	(0.0686)	(0.0949)	(0.0878)	(0.0402)	(0.0326)	(0.0303)
Rubber products	0.7688***	0.9779***	1.0448***	0.0970	0.0256	0.0012
	(0.0743)	(0.1045)	(0.0914)	(0.0510)	(0.0535)	(0.0422)
Plastic products	0.6250***	0.8499***	0.8648***	-0.0468	-0.1024***	-0.1789***
	(0.0769)	(0.0935)	(0.0879)	(0.0526)	(0.0317)	(0.0330)
Petroleum products	0.8578***	1.1433***	1.3004***	0.1860***	0.1910***	0.2567***
	(0.0881)	(0.1057)	(0.0952)	(0.0675)	(0.0615)	(0.0527)
Heavy chemicals	0.7378***	0.9763***	1.0179***	0.0660*	0.0240	-0.0258
	(0.0666)	(0.0955)	(0.0981)	(0.0362)	(0.0352)	(0.0572)
Fertilisers and pesticides	0.8983***	1.2790***	1.2095***	0.2265***	0.3268***	0.1658***
	(0.0790)	(0.1038)	(0.0915)	(0.0536)	(0.0539)	(0.0404)
Paints, varnishes and lacquers	0.8314***	0.9635***	0.9130***	0.1596***	0.0112	-0.1306***
	(0.0802)	(0.1116)	(0.0969)	(0.0571)	(0.0627)	(0.0502)
Drugs and medicines	0.8380***	1.0403***	1.0416***	0.1662***	0.0880**	-0.0021
	(0.0732)	(0.0998)	(0.0898)	(0.0439)	(0.0428)	(0.0389)
Soaps, cosmetics and glycerine	0.7234***	0.9373***	0.9188***	0.0516	-0.0149	-0.1249*
	(0.0772)	(0.1027)	(0.1033)	(0.0510)	(0.0508)	(0.0649)
Miscellaneous chemicals	0.8023***	1.0218***	1.0086***	0.1305***	0.0695***	-0.0351
	(0.0686)	(0.0993)	(0.0950)	(0.0389)	(0.0393)	(0.0502)
Non-metallic mineral products	0.6489***	0.9817***	0.9570***	-0.0228	0.0294	-0.0867***
	(0.0594)	(0.0921)	(0.0857)	(0.0184)	(0.0266)	(0.0273)
Basic metal industry	0.8467***	1.0748***	1.1860***	0.1749***	0.1226***	0.1423***
	(0.0586)	(0.0907)	(0.0846)	(0.0160)	(0.0204)	(0.0240)
Metal products	0.6371***	0.8198***	0.8756***	-0.0347	-0.1325***	-0.1681***
	(0.0592)	(0.0920)	(0.0833)	(0.0227)	(0.0224)	(0.0188)
Machinery for agriculture and food & textile industries	0.7040***	0.9924***	0.8943***	0.0322	0.0401	-0.1494**
	(0.0695)	(0.1004)	(0.1032)	(0.0450)	(0.0444)	(0.0714)
Other non-electrical machinery	0.7548***	1.0200***	1.1339***	0.0830***	0.0678***	0.0902***
	(0.0604)	(0.0925)	(0.0848)	(0.0236)	(0.0226)	(0.0262)

	Wage equation coefficients			Wage premia		
	1983	1993	1999	1983	1993	1999
Electrical industrial machinery	0.9205***	1.0988***	1.0720***	0.2487***	0.1466***	0.0283
	(0.0636)	(0.0992)	(0.0854)	(0.0273)	(0.0407)	(0.0295)
Electrical appliances and electronics	0.7460***	0.9962***	0.9540***	0.0742***	0.0440	-0.0897***
	(0.0632)	(0.0938)	(0.0860)	(0.0283)	(0.0287)	(0.0281)
Sea, rail and motor transport equipment	0.7472***	0.9935***	0.8994***	0.0754***	0.0412*	-0.1443***
	(0.0591)	(0.0908)	(0.0822)	(0.0189)	(0.0239)	(0.0206)
Other transport equipment	0.5207***	0.8724***	0.7995***	-0.1511***	-0.0799**	-0.2442***
	(0.0658)	(0.0972)	(0.0893)	(0.0323)	(0.0398)	(0.0392)
Misc. manufacturing industry	0.6401***	0.9482***	0.9280***	-0.0317	-0.0041	-0.1157***
	(0.0594)	(0.0916)	(0.0838)	(0.0226)	(0.0189)	(0.0215)
Utilities:						
Electricity	0.8548***	1.1685***	1.3900***	0.1830***	0.2163***	0.3463***
	(0.0578)	(0.0921)	(0.0831)	(0.0145)	(0.0179)	(0.0198)
Gas & water supply	0.7255***	1.0361***	1.3337***	0.0537	0.0839**	0.2900***
	(0.0673)	(0.1009)	(0.0906)	(0.0352)	(0.0412)	(0.0387)
Construction	0.6919***	0.9289***	1.0403***	0.0201	-0.0233	-0.0034
	(0.0583)	(0.0913)	(0.0846)	(0.0165)	(0.0213)	(0.0203)
Trade, hotels and restaurants:						
Wholesale and retail trade	0.4659***	0.6989***	0.8261***	-0.2059***	-0.2533***	-0.2176***
	(0.0575)	(0.0897)	(0.0807)	(0.0083)	(0.0088)	(0.0086)
Hotels and restaurants	0.5272***	0.8430***	0.9043***	-0.1446***	-0.1093***	-0.1394***
	(0.0582)	(0.0901)	(0.0815)	(0.0156)	(0.0182)	(0.0175)
Transport, storage and communication:						
Railway transport services	0.7758***	1.0983***	1.3176***	0.1040***	0.1461***	0.2740***
	(0.0565)	(0.0904)	(0.0818)	(0.0086)	(0.0119)	(0.0135)
Other transport services and storage	0.7336***	0.9745***	1.0097***	0.0618***	0.0223**	-0.0340***
	(0.0571)	(0.0906)	(0.0812)	(0.0087)	(0.0106)	(0.0092)
Communication	0.6216***	0.9534***	1.0755***	-0.0502***	0.0012	0.0318
	(0.0588)	(0.0936)	(0.0847)	(0.0180)	(0.0245)	(0.0212)
Services:						
Banking	0.8816***	1.1700***	1.3055***	0.2098***	0.2178***	0.2618***
	(0.0587)	(0.0916)	(0.0816)	(0.0141)	(0.0159)	(0.0162)
Insurance	0.9482***	1.2003***	1.2692***	0.2764***	0.2481***	0.2255***
	(0.0707)	(0.0973)	(0.0897)	(0.0407)	(0.0387)	(0.0396)
Education and research	0.7227***	1.0030***	1.1343***	0.0509***	0.0507***	0.0907***
	(0.0575)	(0.0900)	(0.0821)	(0.0077)	(0.0090)	(0.0104)
Medical and health services	0.6845***	0.9992***	1.1741***	0.0127	0.0470**	0.1304***
	(0.0581)	(0.0933)	(0.0830)	(0.0153)	(0.0199)	(0.0200)
Other services [a]	0.4910***	0.7480***	0.8375***	-0.1808***	-0.2042***	-0.2062***
	(0.0574)	(0.0904)	(0.0826)	(0.0111)	(0.0114)	(0.0161)
Public administration	0.7195***	1.0142***	1.2175***	0.0477***	0.0619***	0.1738***
	(0.0566)	(0.0901)	(0.0811)	(0.0043)	(0.0053)	(0.0057)
Employment-weighted mean standard deviation [b]				0.1338	0.1413	0.1844

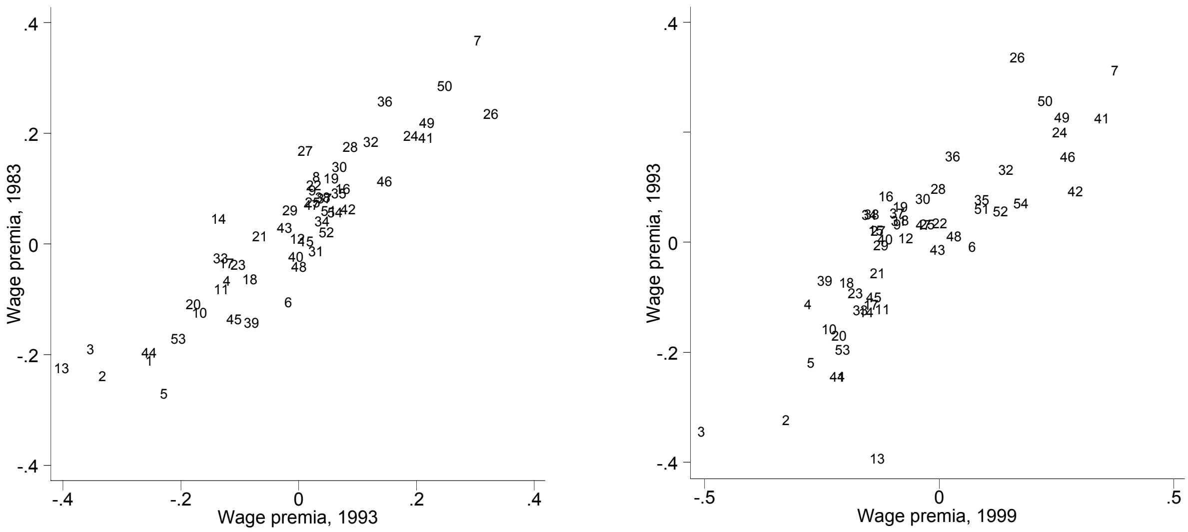
Notes: 1\ * significant at 10%; ** significant at 5%; *** significant at 1%. 2\ Standard errors are in parentheses computed using the procedure suggested by Haisken-DeNew and Schmidt (1997). a\ Other services comprises legal, business, personal, social, sanitary and community services. b\ This is a summary measure of the overall variability in wages across industries computed using the procedure suggested by Haisken-DeNew and Schmidt (1997).

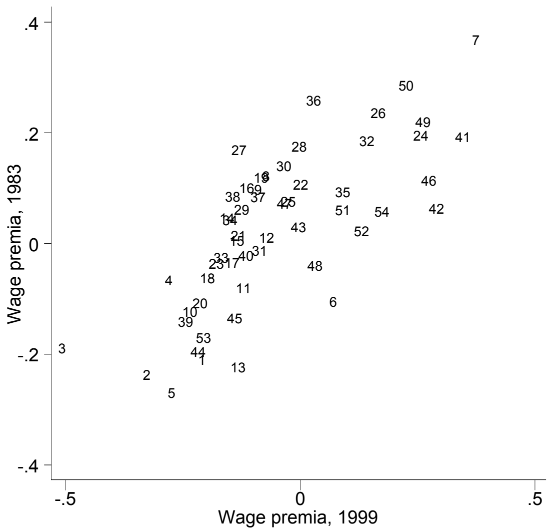
Table A3: Summary statistics by quantiles of tariff distribution, 1983

	Low tariff		Medium tariff		High tariff	
	Mean	**S.D.**	**Mean**	**S.D.**	**Mean**	**S.D.**
Average tariff rate	0.8213	0.1809	1.0463	0.0445	1.2977	0.2283
Wage premia	0.0466	0.1349	0.0320	0.1280	0.0260	0.0845
Share of female workers	0.1417	0.1711	0.1515	0.1374	0.1239	0.0813
Share of casual workers	0.1770	0.1449	0.1810	0.1118	0.1519	0.0613
Share of skilled workers	0.2964	0.1983	0.2771	0.1791	0.2035	0.1292
Pro-worker state	0.4377	0.1745	0.3345	0.1429	0.3676	0.2027
Anti-worker state	0.3658	0.1671	0.4166	0.1126	0.3696	0.1538
Neutral state	0.1965	0.0933	0.2489	0.1368	0.2628	0.0781
License dummy	1	0	1	0	1	0
Public dummy	0.3636	0.5045	0.3000	0.4830	0.0909	0.3015
Union density	0.2388	0.1762	0.2442	0.1897	0.1761	0.1055
Average establishment size	168.4065	259.5839	80.9690	38.0550	59.9211	28.9501
Capital-output ratio	0.2803	0.1606	0.3366	0.2527	0.2832	0.1459
Capital-labour ratio	0.7120	1.0949	0.4652	0.4073	0.4116	0.2768
Share of registered manufacturing	73.1990	15.6554	73.7250	14.7567	60.0343	31.7988
Real imports	42,352	61,095	38,105	39,475	20,128	48,206
Real exports	28,337	31,329	28,800	37,318	17,831	30,431
Industry import share	0.0269	0.0388	0.0242	0.0251	0.0128	0.0306
Industry export share	0.0267	0.0295	0.0271	0.0352	0.0168	0.0287
Consumer goods	0.4545	0.5222	0.4000	0.5164	0.6364	0.5045
Capital goods	0.2727	0.4671	0.2000	0.4216	0.0909	0.3015
Intermediate goods	0.1818	0.4045	0.3000	0.4830	0.1818	0.4045
Basic goods	0.0909	0.3015	0.1000	0.3162	0.0909	0.3015
Number of observations	11		10		11	

Notes: The tariff distribution for manufacturing industries has been divided into 3 quantiles – the low tariff group has industries with tariffs less than 1.01 and the high tariff group comprises of industries with tariffs greater than 1.14 in 1983.

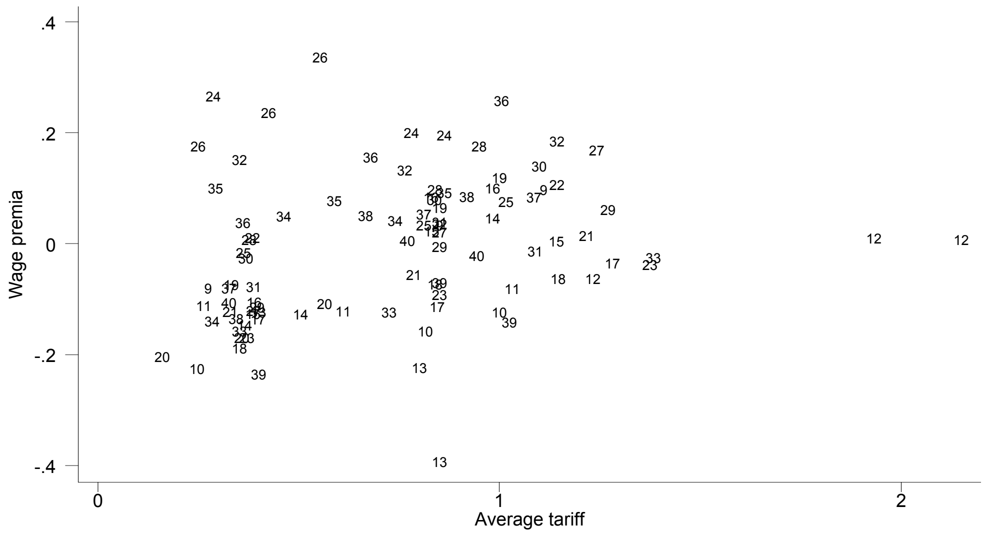
Figure A1: Wage premia: 1983 to 1999





Note: The numbers represent the industry: Industries 1-6 are agriculture and allied activities, 7-8 are fuel extraction and mining, 9-21 are light manufacturing, 22-40 are heavy manufacturing, 41-42 are utilities, 43-45 are construction, trade and hotels, 46-48 are transport, storage and communications, 49-53 are services and 54 is public administration (see Table A1 on summary statistics for details).

Figure A2: Wage premia and tariffs, 1983 to 1999.



Note: The numbers represent the industry code: Industries 9-21 are light manufacturing, 22-40 are heavy manufacturing, 12 is the beverages industry (see Table A1 in the Appendix for details).