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**ECO 2007/50**

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ISSN 1725-6704



Printed in Italy
European University Institute
Badia Fiesolana
I – 50014 San Domenico di Fiesole (FI)
Italy

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# Do Unions Matter?
# Trade Reform and Manufacturing Wages in South Africa *

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November 7, 2007

**Abstract**

**The effect of *nominal* tariff cuts on industry wage differentials has been the subject of a number of recent empirical studies. In this paper we investigate the latter relationship with respect to the South African trade reform experience using labor force data for the period from 1995 to 2004. Our study extends on the existing literature in two respects: firstly, we control for the potential effect of labor market institutions, such as collective bargaining power, in assessing the relationship between tariffs and industry wages. Secondly, we account for general equilibrium effects by controlling for the impact of changes in *effective* tariffs rates. We find that on the one hand, only wages in industries with levels of unionization beyond a certain threshold were adversely affected by tariff cuts. This negative effect is exacerbated by the extent of sectoral union power. The reported large magnitudes of the tariff impact on wages is in line with the considerably high markups documented for South Africa. On the other hand we find some evidence suggesting that wages in industries with union power below the threshold were positively affected by the tariff cuts.**

KEYWORDS: **Trade Reform, Tariffs, Manufacturing Wages, Trade Unions, South Africa**
JEL CLASSIFICATIONS: **F14, F16, O55**

*Helpful comments from Morten Ravn are gratefully acknowledged. I also thank Lawrence Edwards for making the detailed tariff data available and UC Berkeley Development Lunch for their useful feedback.

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# 1 Introduction

In this paper we investigate the impact of the tariff reductions in South Africa (SA) on worker wages in the manufacturing sector during a period of intensive trade reform. From 1995 to 2004 the country witnessed considerable tariff cuts. Average Nominal Tariff Rates (NTR) and average Effective Tariff Rates (ETR) in the manufacturing sector decreased from pre-reform levels of 20% and 48% respectively in 1993 to 7% and 13% in 2004, with large dispersions across the different sub-industries. This paper extends on the existing literature that investigates the effect of trade liberalization on wages premiums along two lines. Firstly, this work is the first to control for the effect of collective bargaining on wages when examining the trade-wage relationship. We hypothesis an asymmetric effect of tariff reductions on industry wages due to inter-industry differences in union bargaining power. Accounting for collective bargaining is particularly relevant in the case of SA given the central role played by trade unions. We support our hypothesis by outlining theoretical models that: (1) suggest a correlation between trade policy and industry bargaining power, and (2) predict an asymmetric industry response to trade openness based on differences in sectoral bargaining power. Secondly, in addition to using NTR as the proxy to measure changes in trade policy, our paper provides more evidence on general equilibrium effects as we control for changes in ETR. The latter measures the total effect of protection on output in addition to the cost raising effect of protection on intermediate inputs, by that reflecting the intermediate input linkages across industries.

Combining, both, household and labor force data for SA from 1995 to 2004 we exploit the variation in tariff rates, union power and wages, across industries and over time, to investigate the asymmetric effect of tariff cuts on wages attributed to heterogeneities in sectoral bargaining power. Our results suggest that the negative impact of tariff cuts on wage premiums is only conditional on the extent of the industry's union bargaining power, the latter proxied by sectoral union density. There is a threshold level of union density after which tariff reductions lead to significant wage declines. An increase in union bargaining power beyond this threshold is associated with a an even larger negative impact of tariff cuts on wages. We find that the magnitude of the negative impact of tariff cuts on wages is considerably large when compared to findings in the literature for other countries. This large effect is not surprising in light of the significantly large markups documented for SA, which are twice those reported for the US and higher than those documented for OECD countries (Fedderke et al 2006, Aghion et al 2006). Accordingly our results are consistent with the notion that wage adjustment is larger where

wages exceed competitive market levels. Furthermore, we find some evidence suggesting that industries with union power below the threshold witnessed a rise in wages as displayed by the positive coefficient on the interaction of tariffs and union density and by the negative coefficient on the tariff variable. Particularly for SA it is important to note the well documented political reasoning behind unionization which is primarily motivated by historical racial concerns as opposed to the commonly known economic objectives of trade unions (Azam & Rospabe 2007). In light of the latter and given our focus on assessing the role of industry-level unionization, as opposed to worker union membership status, in the trade-wage relationship, and additionally given our control for workers' race in our regressions, we treat industry-level unionization as exogenous to workers' wages.

In this work we also account for concerns with regards to the endogeneity of trade protection as suggested by theories of political economy predicting that product and labor market concerns are likely to be the basis for the chosen trade polices. We show that our results are not driven by this endogeneity bias. Firstly, we argue that the structure of the tariff schedule reveals limited industry selectivity tariff cuts. This is demonstrated in the tariff data which shows that industries with higher initial tariff levels proportionately witnessed the larger tariff reductions. This feature of the tariff series emphasizes that industry lobbying were restricted as all tariffs approached commonly low levels. Albeit, we still account for the plausible endogeneity of trade policy by: (1) considering the effect of one period lag tariffs in our estimation, (2) controlling for unobserved industry fixed characteristics by including industry fixed effects, and (3) we use pre-sample (and pre-reform) 1993 tariff levels and their interaction with the foreign exchange rate, or alternatively with gold prices, as instruments for the *changes* in tariff rates.

Gaston and Trefler (1994) is one of the first attempts to explore the direct link between tariffs and industry wage premiums. Using data for the US manufacturing sector in 1983 they find a negative effect of NTR on wages. With respect to developing countries Currie & Harrison (1997) and Revenga (1997) are the earliest studies addressing this same relationship. Focusing on Morocco, Currie & Harrison (1997) find that NTR cuts had an insignificant effect on wages, while Revenga (1997) shows that Mexican NTR reductions were associated with significantly lower wages. In more recent studies Goldberg et al (2005) and Pavcnik et al. (2004) examine the evolution of industry wage premiums following the cuts in NTR witnessed for Colombia and Brazil, respectively. Controlling for industry specific effects, Goldberg et al. find that industries facing higher tariff cuts experienced larger reductions in wages. Pavcnik et al. find instead tariffs to be insignificant in explaining period variations in wage premiums in Brazil. Similar work is

conducted on India (Mishra and Kumar, 2005; Vasudeva-Dutta, 2004), Mexico (Feliciano, 2001), the Philippines (Hasan and Chen, 2004) and Poland (Goh and Javorcik, 2004). With respect to India, results from the former work support that trade liberalization induced higher firm productivity leading to higher wage premiums. Vasudeva-Dutta (2004) using a different dataset finds the opposite. Results on Mexico and the Philippines are similar to those reported for the Brazilian experience suggesting the insignificant tariff-wage relationship. As for Poland, findings from Goh and Javorcik (2004) show that NTR reductions are associated with higher wages.

Some of the recent empirical trade literature focuses on the role of industry bargaining power in determining the impact of changes in trade policy on labor market outcomes. However, these studies refer to developed economies and only make use of trade policy outcomes, as opposed to policy instruments, as the proxy to measure trade openness. Macpherson and Steward (1990) study the impact of changes in imports on unionized versus non-unionized wages in the US. They find that increases in imports were associated with lower wages and that a high level of industry unionization decreases the negative impact of increased import competition on wages. Freeman and Katz (1991) compare the responsiveness of US wages to sales shocks, resulting from international trade across industries with different levels of unionization. They find that wages were more responsive to sales shocks in the highly unionized industries. A more recent study by Griffith et al. (2005) analyze whether the effect of product market competition on real wages depends on the level of collective bargaining power. Using data on OECD countries they find that lower mark-ups, which reflect the increase in foreign competition, led to higher real wages. Yet their results show that the interaction between the union density variable and industry mark-ups was positive and significant, suggesting that the positive effect of reduced mark-ups on wages decreases in the presence of higher bargaining power.

This paper is divided into seven sections. In the next section we introduce SA's trade policy and unionization background. In Section 3 we outline the relevant theoretical models. Section 4 describes both the labor force and trade data on hand. Section 5 presents the empirical methodology. Section 6 discusses our results. And finally Section 7 concludes.

# 2 Country Background

## 2.1 Trade History

Up until 1970s SA was firmly oriented towards Import Substitution Industrialization . The latter consisted of a wide-ranging system of Quantitative Restrictions (QR) as opposed to tariff-based

protection. The first shift away from this trade regime came in 1972 with the relaxation of QR and the introduction of an Export Development Assistance scheme, however the overall trade policy remained protectionist. In the mid 1980s SA faced balance of payment pressures arising from a debt crisis which led to the implementation of import surcharges which offset the effects of the QR relaxation that continued into the 1990s and were completed by 1994. Moreover in 1990 the General Export Incentive Scheme (GEIS) was introduced granting subsidy to exporters based on their export value.[1]

In April 1994, the first post-apartheid government led by the African National Congress (ANC) party was democratically elected. The ANC party was strongly supported by Congress of South African Trade Unions (COSATU), the largest labor federation in SA. This ANC-COSATU alliance explains why many of the trade union leaders became prominent members of the new government. Simultaneously, multilateral trade reform was initiated in 1994 with the WTO agreeing on the phase-down tariff plan offered by SA in the GATT/WTO Uruguay Round. By signing the latter agreement the country committed to significantly reducing tariff rates. Consequently during the period from January $1^{st}$ 1995 onwards SA experienced considerable cuts in protection rates. Furthermore, the same period was associated with the decision to phase out of GEIS where by 1997 the export subsidy provided under the GEIS was terminated.

Investigating trade liberalization episodes across developing countries, Michaely et al. (1991) find that radical political changes towards more stable regimes are usually accompanied by adopting free trade policies. Differently, for SA it is believed that the support of the country's Industrial Development Corporation (IDC) for trade policy reform was a result of the anticipated change in the regime. The IDC believed that committing to the trade reform would limit the intervention of the ANC-COSATU led government in the country's industrial development and would restrict the power of trade unions in determining the course of economic development in the post-apartheid era. Contrary to expectations, the ANC-COSATU alliance coming to power in 1994, proved to hold an anti-protectionist stand as they actively conformed to the stipulated tariff reduction schedule. As documented in Bell (1997) the new government's liberal position was triggered by their plan of reducing consumer prices and raising industrial efficiencies by curbing domestic monopoly power and the control of conglomerates that had vested interests in the prevalent protectionist policies. [2]

[1] See Bell 1997 for a more detailed description of SA Trade Policy

[2] A thorough characterization of the tariff structure will be presented in Section 4.

## 2.2 Unionization History

A prominent feature of the South African labor market is its high rate of unionization. Based on our dataset, the economy wide unionization rate averaged to 35% in the period from 1995 to 2004 (of which 70% black & 64% male). This figure is fairly high when compared to the average rates in developing countries estimated at 20%, and is more in line with figures for developed countries estimated to average to 40% for OECD countries. Another feature of the South African trade unions that resembles developed countries concerns union membership wage premiums which are estimated to range from 10% to 24%, comparable to 10% reported for UK and 15% for the US (Aidt et al, 2002).

Collective bargaining in SA takes place at, both, a centralized and a decentralized level. The centralized level operates under Bargaining Councils (BC).[3] BC agreements are established when registered employer associations voluntarily agree to bargain with registered trade unions. The agreements cover a specific industry, occupation, geographic location (Bendix 1989) and operate under *ergo omnes* rules [4]. BC agreements can not be established unless they represent the majority of workers in the agreement's area of specification and would represent the majority of employers. Differently, decentralized bargaining takes place at the firm level (Azam et al, 2007). The Congress of South African Trade Unions (COSATU), established in 1985, is by far the biggest and most effective of the country's three main labor federations. Based on the COSTAU $8^{th}$ Congress Organizational Review, average union rates declined since 1997 from a rate of 36% to 32% in 2003. This erosion of membership is believed to be due to the massive layoffs as unemployment rates increased from 22% in 1997 to reach a maximum of 30.5% in 2003 (Source: IMF).

To capture the centralized and decentralized collective bargaining effect on wages it would be ideal to have information on the union status of the establishment to which the worker belongs and whether workers in the establishment are covered by a BC agreement. Such data is not available. In place we use worker union membership status, as provided by the household surveys, to derive our proxy for industry-level union bargaining power which is based on union density in an industry. With regards the BC centralized effect on wages, we capture the latter by the industry, occupational and regional control dummies.

With regards to the endogeneity of industry-level unionization to workers wages, we argue

[3]Prior to the new labor relations act of 1995 BC were known as Industrial Councils

[4]Meaning that wage agreements extend to all workers covered by the area of the agreement whether unionized or not. (Butcher & Rouse 2001)

that given the particularity of SA and the race motivated history underlying unionization, we dismiss the latter as a concern in our estimations.[5] The latter concern is further reduced in light of our focus on assessing the role of industry-level unionization, as opposed to worker union membership status, in the trade-wage relationship.

# 3 Theoretical Background

In this section we start by outlining the theoretical argument supporting the correlation between trade policy and industry union power. We rely on this theoretical framework to emphasize the potential omitted variable bias stemming from not controlling for industry unionization as the latter is likely to affect industry wages independently from the impact of tariffs. Given the focus of this paper we do not argue in favor of a particular causal relationship between trade policy and union power. We merely highlight the presence of a strong correlation between the latter variables. Secondly, we outline theories predicting the differential effect of trade policy on wages attributed to heterogeneities in the level of industry bargaining power which supports our hypothesis of a significant role for the interaction between tariff reductions and union power.

## 3.1 The Correlation Between Trade Policy and Industry Union Power

Union Density, defined as the number of union employment in a specific industry over total employment of that industry, is the popular proxy for union bargain power. Hirsch & Schumacher (2002) discuss two primary channels that affect this ratio. The *first* relates to the changes in union wage premiums. In the context of free trade policy, workers' sentiment towards joining unions may change due to a perceived decline in union bargaining power resulting from trade induced reductions in industry rents. Alternatively, employers opposition towards employing unionized workers can adversely affect the ratio given the negative effect of unions on profits (Addison & Hirsch 1989). Additionally, the increasing threat of international capital mobility, induced by the rising measures of globalization, raises employers anti-unionization sentiment as it becomes easier to relocate businesses to new markets with more flexible labor regulations (Farber & Western 2000). The *second* channel takes place through alterations in the labor market structure due to shifts in labor demand or supply associated with changes in trade policy. These can take the form of changes in industry, occupation, demographic structure, or regional location of jobs. In the framework of trade reform, trade openness may increase employment of labor segments that are traditionally less likely to be unionized such as skilled

[5]Refer to Azam & Rospabe (2007) for a more thorough discussion on the racial history of trade unions in SA.

labor or female workers. For instance, an increase in skill employment can be stimulated by trade induced Skill Biased Technological Changes (SBTC), whereas rises in female participation can be due to labor reforms which are likely to go hand in hand with trade reforms.

Arguing the reverse relationship where collective bargaining affects trade policy decisions, Rama (1997), motivated by the strikes in Zambia, Nigeria and Venezuela that led to reversing reform programs, develops a model based on a two-stage game between the government & organized labor to determine the level of product market distortions proxied by import tariff rates. The model predicts that organized labor are more willing to cooperate (as opposed to strike) the larger is the percentage of the labor force enrolled in trade unions. A similar result is presented in Rama & Tabellini (1998). The underlying logic is that workers gain from higher import tariffs in their own sector but as consumers they are better off if tariffs for all other sectors are low. Accordingly if trade unions are narrowly defined, their members benefit greatly from tariff barriers and wage differentials while imposing only a light burden on each individual consumer. But as the number of beneficiaries from the distortions increases, the burden gets heavier leading organized workers to internalize (as consumers) part of the resulting efficiency losses. This follows from the logic of *Virtues of Corporatism.*

### 3.2 Models of Trade, Collective Bargaining and Industry Wages

Under the assumption of perfect competition the conventional Heckscher-Ohlin Model predicts a heterogenous effect of trade policy on industry wages on the basis of sectoral differences in worker specific characteristics. Alternatively, under the Specific Factor Model winners and losers of trade are directly identified by their industry affiliation. Abstracting from a world of perfect competition, theories of imperfect markets such as Efficiency Wage and Rent-Sharing allow for wide inter-industry wage dispersions offering alternative channels through which free trade policies can impact industry returns.

In the presence of collective bargaining, rent-sharing models constitute the relevant theoretical framework. Monopolistic Competition in the product market determines size of industry rents. Bargaining power in the Labor Market determines the distribution of rents between workers and the firm. In an efficient bargaining framework, the higher an industry's union power the more it is able to capture portions of rents in the form of higher wages for their unionized workers. Freeman and Medoff (1981) and Hirsch and Neufeld (1987) show that the ability of unions to negotiate higher wages for unionized workers is a positive function of the degree of unionization. This positive impact of unions may further extend to also affect wage of non-

unionized worker if the *threat* or *demand* effect offset the *supply* effect (Hirsch and Addison, 1986). Consequently, one would expect a higher wage mark-up in industries with stronger union representations.

Furthermore, trade union theory suggests that unions are able to capture excess rent in the presence of protection (Lawrence and Lawrence, 1985). With the increase in foreign competition induced by free trade policies and consequently the congestion of industry rents, unions face a trade-off decision between lower employment and lower wages and thus may bargain for employment guarantees at the expense of accepting wage concessions. This notion is more formally presented in Freeman & Katz (1991). Moreover, Abowd and Lemieux (1991) show that if international competition adversely affects total quasi rents[6] available to both the employer and unions, then either wage settlement or union employment will be negatively affected. Interestingly, Grossman (1984) presents a model in which the opposite may occur. He finds that majority voting in unions coupled with a seniority system for layoffs and rehires lead to sticky wages which do not adjust to intensified international competition, thereby emphasizing the important role of the seniority structure of unions in determining union wage response to trade reform.

In light of the above discussion we believe that there is ample reason in support of accounting for industrial collective bargaining when investigating wage responsiveness to changes in trade policy.

# 4 Data

## 4.1 Labor Force Data

In this paper we use repeated cross section labor force data from two independent surveys provided by the South African Data Archive (SADA). The first is the annual October Household Survey (OHS) which runs from 1993 to 1999. We exclude three rounds ; the 1993 OHS being a pilot survey with insufficient information on workers' industry affiliation, the 1994 OHS whose information we find unreliable[7] and the 1996 OHS due to the absence of the wage remuneration point values as wages are reported in income intervals. The second dataset we use is from the Labor Force Survey (LFS) which starts in 2000. The latter is a twice-yearly rotating panel

[6]The difference between net revenues and cost of employment.

[7]We find that the wage distribution and sample composition for 1994 if considerably different from both: that of 1993 (using the South African Labor and Development Research Unit (SALDRU) dataset also known as the Living Standard Life Survey (LSMS)) and that of the proceeding years. We do not use the SALDRU 1993 data in our analysis as it does not provide worker sub-industry affiliation within the manufacturing industry.

conducted in February and September of each year. In our study we use the September wave. Matching both datasets we are able to cover nine years of the trade reform period, from 1995 to 2004 (excluding 1996).

Our dataset covers all of South Africa's nine provinces and in this respect is representative of the country's labor market. We restrict our sample to workers in the manufacturing sector of age 15 to 65 with reported monthly wages between 100 and 30,000 Rands. Given the focus of the paper on the relationship between tariffs, wages and union power, we exclude self employed workers. Consequently we are left with a sample of 17,329 observations which vary from a maximum of 2,965 in 1995 to a minimum of 1,036 for 1998. With regards to our dependent variable, the wage remuneration question 'what is total salary/pay (including overtime) at main job (before deductions)' does not change throughout both surveys. We are unable to report hourly wages given the lack of information on the number of hours usually worked in the 1995 survey accordingly we are forced to use monthly wages.[8] We control for workers' age, household size and for whether the individual is a male, head of household and single. We break the sample into three racial groups black, others and white the latter being the regression Omitted Category (OC). Controlling for workers' race is particularly important in this study given the racial motivation underlying union membership in SA. Using information on the highest completed level of education we construct five education classifications: no schooling (OC), not completed primary schooling, completed primary schooling, completed secondary schooling and completed a university degree or above. We further control for geographic location by distinguishing 9 provinces. We define 9 occupational categories, and we report only 10 categories for workers' industry affiliation for two reason: firstly, due to the lack of more detailed industry affiliation information for the 1995 OHS. Secondly, to ensure that we have a fair number of observations under each industry at every point in time. Table 1 shows the breakdown of the occupation and industry categories.

Examining the evolution of wages during the period, there is a general decline in both the mean and the median of log monthly wages, particularly during the first three years of the period. This decline is further confirmed by the leftward shift in the wage distribution (see Figure 1). We use the survey information on workers' union membership status to construct our industry union density ($UD$) variable computed as the percentage of union members in a particular industry of the total workers employed in this industry. The union status question 'is ... a member of

[8]Excluding 1995 from the estimation is costly given that tariff rates were substantially reduced between 1994 and 1997

a trade union?' does not change through out the surveys ensuring the time consistency of our union power estimate. The yearly average union density across all 10 manufacturing industries exhibits a declining trend (Figure 2). Based on our data, the period average rate of unionization in the manufacturing sector is estimated at 43% (of which: 66% male, 68% African). We find highest rates of 55% and 47% for TRANS and TEX sectors respectively, and a lowest of 27% for the NoMet sector.[9]

### 4.2 Tariff and Industry Level Data

Regarding the tariff data, we use NTR and ETR provided in Edwards (2005).[10] The original data series is at the SIC-3 digit level. For the reason previously noted we aggregate our tariff data to 10 industries using the 1993 pre-sample imports as weights. Table 2 summarizes the tariff data. Average NTR decreased from 19.62% in 1993 to 6.68% in 2004 while ETR declined from 47.75% to 12.79%. Notably the first three years of the reform witnessed the highest reductions in protection rates where by 1997 average tariffs had reached less than 50% of their 1993 pre-reform value. Moreover the decline in yearly standard deviation indicates the decreasing dispersion in the cross industry levels of protection as they approach commonly low levels. Figures 3-A and 3-B plot the pre-sample tariff rates in 1993 against the change in tariffs between 1993 and 2004.[11] The downward sloping graph indicates that industries with initially high levels of protection experienced more severe liberalization measures. This feature of the data implies that industry selectivity reduction or lobbying were to some extent limited during the reform.

We use data on capital, employment, exports and imports from the South African Standardized Industry Database - SASID (Quantec Research, 2006) to control for industry characteristics. We aggregate the SIC-3 digit data to our 10 industry categorizations to match our labor force data. GDP figures at the 10 industry classification are obtained from the GDP annual reports provided by Statistics South Africa's (STATSSA) . Using this data we construct variables that measure trade flows such as: export intensity and import penetration ratios.[12] In our regressions we also control for changes in international prices. There is no available direct measure for the international prices faced by the South African consumers, accordingly and in line with Currie & Harrison (1997) we use the US export price index series calculated by the Division of International Prices at the US Bureau of Labor Statistics. We roughly create a concordance

[9] Refer to Table 1 for abbreviations.

[10] We use the tariff data computed using scheduled tariff rates and inclusive of surcharges. A detailed description of the methodology adopted for the tariff calculations is provided in the paper.

[11] We compute the tariff changes, $\widetilde{\Delta}$, as noted in Table 2.

[12] Export Intensity = $Exports/GDP$, Import Penetration = $Imports/(GDP + Imports - Exports)$.

between the US export price series which is provided at the Harmonized System (HS) classification and our industry categorization. We also use South Africa's export price index provided by *Quantech* as another proxy for international prices. Under the assumption that export prices of South African producers are solely determined by the international market, we consider that the latter proxy is exogenous to worker wages.

# 5 Empirical Methodology

In this section we outline the two empirical techniques implemented in this paper to estimate the effect of reductions in NTR and ETR on wages, controlling for the role of collective bargaining. The first procedures entails estimating an expanded Mincerian wage regression. With regards to the second technique, we adopt a two stage estimation framework that enables estimating the latter relationship in First-Difference which accordingly allows using pre-reform tariff levels in 1993 as an instrument for tariff *changes*.

## 5.1 One-Stage Estimation Procedure

We estimate an expanded Mincerian wage regression, equation (1), in which we regress real log monthly wages, $w_j$[13], on a matrix of worker characteristics, $Q$[14], NTR and ETR ($\tau$), union density ratio ($UD$), the interaction of the latter ($\tau * UD$), and a matrix of other industry and trade related controls ($H$). We include time dummies ($TD$) to absorb macroeconomic shocks and 9 industry dummies ($ID$) in order to control for time invariant unobserved industry specific heterogeneities affecting simultaneously wages and tariffs. Moreover following Currie & Harrison (1997), where domestic prices are given by $P = P^*(1+\tau)$, we control for the international price, $P^*$, to account for the fact that changes in international prices may dampen the effect of tariff cuts on wages.

$$Ln(w_{ijt}) = \beta_Q Q_{ijt} + \beta_\tau \tau_{jt} + \beta_{UD} UD_{jt} + \beta_{(\tau * UD)}(\tau_{jt} * UD_{jt}) + \beta_H H_{jt} + \beta_{ID} ID_{ijt} + \beta_{TD} TD_{ijt} + \epsilon_{ijt} \quad (1)$$

The total effect of tariffs on wages is given by $\beta_\tau + \beta_{(\tau * UD)} UD$. On the one hand, if our hypothesis is true and tariff reductions are associated with larger wage declines for the highly unionized sectors, then we expect $\beta_{(\tau * UD)}$ to be positive. On the other hand, $\beta\tau$ captures the

[13] Nominal monthly wages are deflated using the Consumer Price Index from STATSSA.

[14] Age, age squared, gender, race, position in the household, marital status, provincial location, education and occupation.

total effect of tariffs on wages for the non-unionized sectors. A negative coefficient implies that tariff reductions were associated with a rise in wages in those sectors.

### 5.2 Two-Stage Estimation Procedure

Theories of political economy predict that product and labor market concerns are likely to be the basis of the implemented trade polices suggesting the endogeneity of protection. As noted in the introduction, we address this issue by: (1) we argue that the structure of the tariff schedule reveals limited industry selectivity tariff reductions, (2) we consider the effect of one period lag tariffs in our estimation, (3) we control for unobserved industry fixed characteristics by including industry fixed effects. Finally, and as will be more thoroughly discussed in Section 6.2, pre-sample 1993 tariff levels and their interaction with real effective exchange rate, or alternatively gold prices, comprise plausible instruments for tariff *changes*. Notably, that we can not estimate equation (1) in First-Difference as the data on hand is a repeated cross section as opposed to a panel dataset of workers. In this respect we adopt the Two-Stage estimation framework commonly used in the trade and labor literature on industry premium (Krueger and Summers 1988, Katz and Summers 1989, Gaston and Trefler 1994, Goldberg et al. 2004). This technique enables constructing a panel of estimated industry wage premiums from the first stage of the procedure which allows an estimation in First-Difference in the second stage. The latter is important for our purpose of instrumenting for *changes* in tariffs. In the first-stage we estimate equation (2). The coefficients on industry dummies, $\beta_{ID}$, capture the part of the variation in wages that is only explained by industry affiliation. We adopt the Haisken-DeNew and Schmidt (1997) two-step Restricted Least Square ($RLS$) procedure which involves including *all* 10 industry dummies and estimating equation (2) subject to the constraint (3), where $n_j$ is the vector of each industry's employment share. By using the latter technique to normalize the industry wage differentials we avoid the dummy variable trap which requires an omitted-control group. Hence industry coefficients can be interpreted as the proportional difference between the log monthly wage for workers in a given industry, $j$, and the employment share weighted average log monthly wage of workers (with the same observable characteristics) in all industries. Moreover, this method provides correct standard errors for the estimated $\beta_{ID}$ coefficients.[15]

[15]Krueger & Summers (1988) used a two step re-normalization to express industry differentials as deviation from an employment-weighted average. They approximate the standard errors (SE) of the reorganized coefficients by the SE of the original regression coefficients. With regards to the omitted control industry, they use the SE of the constant term. Haisken-DeNew & Schmidt (1997) show that this procedure affects inference by overstating the SE of the re-normalized coefficients. They thus present the aforementioned one step RLS procedure for calculating the exact SE of the re-normalized coefficients.

Given that our sample is composed of independent cross section data, we estimate equation (2) separately for each of the years in the sample.

$$Ln(w_{ijt}) = \beta_Q Q_{ijt} + \beta_{ID} ID_{ijt} + \epsilon_{ijt} \tag{2}$$

$$\sum_{j=1}^{10} n_{jt} * \beta_{ID,jt} = 0 \tag{3}$$

In the second-stage estimation we pool the yearly estimated industry wage premiums, $\beta_{ID}$, and regress them on tariff ratios, industry union density ratio, the interaction of the latter, and a matrix of other industry and trade related controls. We control for time, fixed effects and $P^*$ and estimate (4) in First-Difference:

$$\Delta\beta_{ID,jt} = \beta_\tau \Delta\tau_{jt} + \beta_{UD}\Delta UD_{jt} + \beta_{(\tau * UD)}\Delta(\tau_{jt} * UD_{jt}) + \beta_H \Delta H_{jt} + \nu_{jt} \tag{4}$$

Given that the dependent variable in the second-stage is the estimated coefficients from the first-stage wage equation, we use the inverse of $(\sigma^2_{\beta_{j,t}} + \sigma^2_{\beta_{j,t-1}})$ as weights.[16] This technique allows us to assign lower weights to industries with higher variances. Accounting for general forms of heteroscedasticity and serial correlation in the error term, we compute robust (Huber-White) standard errors. Under this Two-Stage methodology we can instrument for the *changes* in tariff rates using initial reform tariff levels and their interaction with foreign exchange rate and with gold prices.

# 6 Estimation Results

## 6.1 First-Stage Results for the Two-Stage Procedure

Table 3 shows the estimation results from wage equation (2) for each year of the sample. Individual characteristics such as older age, being male, and a household head are associated with higher wages. As one would expect, white workers earn the most and black the least. With regards to the education variables, it proves to be significant for most years and across most of the different schooling classifications. Figure 4 plots the coefficients of the four educational dummies with respect to the omitted category being the lowest educational group (no schooling). As demonstrated by the graph, workers belonging to the highest educational category (completed university degree or above) earn substantially higher returns compared to the other groups.

[16] The variance of the dependent variable in the FD estimation of equation (2), $\text{Var}(\beta_{j,t} - \beta_{j,t-1})$, is given by $\sigma^2_{\beta_{j,t}} + \sigma^2_{\beta_{j,t-1}}$ where $\beta_{j,t}$ & $\beta_{j,t-1}$ are assumed to be independent.

The period average parameter for the top educational group is estimated at 1.04 compared to parameter estimates of 0.57, 0.29 and 0.10 for workers who completed secondary schooling, primary schooling and those who did not complete primary schooling, respectively. Yet, unlike the Colombian experience (Goldberg 2005), South Africa's trade reform period was not associated with substantial increases in returns to workers with college education. Accounting for work characteristics by controlling for occupation and industry affiliation considerably increases the explanatory power of the model as the period average R-squared increases from 34.8% to 47.8%. As displayed in Table 3, the coefficients on the occupation dummies (OD) are significant and indicate higher premiums to the more skilled occupations.

Regarding the estimated coefficients on the 10 industry dummies (ID), most of the parameters are significantly different from zero. The bottom row of Table 3 reports the standard deviations ($SD$) of the estimated industry wage premiums weighted by industry employment and adjusted for least square sampling error, as computed by the formula:

$$SD(\beta) = \sqrt{\frac{\sum_j n_j(\beta_j - \beta_j^*)^2 - \sum_j n_j \sigma_{\beta_j}^2}{\sum_j n_j}} \tag{5}$$

where $n_j$ is a vector of employment share in each industry, $\beta_j$ and $\sigma_{\beta_j}$ are the estimated industry premiums and their standard error, and $\beta_j^*$ is the mean of $\beta_j$ weighted by $n_j$ (see Moll 1993). The estimated variation ranges from 6.5% to 16% suggesting that movement across industries has a considerable impact on wages. Notably the period is characterized by a rise in the yearly dispersions across industry premiums as indicated by the upward trending yearly standard deviations (see Figure 5). We compare the weighted and adjusted year to year correlations of the industry premiums given by:

$$Correlation = \frac{\sum_j N_j[(\beta_{j,t} - \beta_{j,t}^*)(\beta_{j,t-1} - \beta_{j,t-1}^*) - \sigma_{\beta_{j,t}}\sigma_{\beta_{j,t-1}}]}{\sqrt{\sum_j N_j[(\beta_{j,t} - \beta_{j,t}^*)^2 - \sigma_{\beta_{j,t}}^2] * \sum_j N_j[(\beta_{j,t-1} - \beta_{j,t-1}^*)^2 - \sigma_{\beta_{j,t-1}}^2]}} \tag{6}$$

where $N_j$ = the geometric mean= $(n_{j,t}n_{j,t-1})^{0.5}$ . This measure of persistence indicates that the pattern of cross-industry wage premiums was relatively unstable in the early past of the sample. This non-persistent structure of industry wage differentials suggests the plausible impact of trade reform on wages which was more intense in the start of the period. The correlation figures pick up to higher levels in the last five years understudy during which the reform took a smoother pace.

### 6.2 The Mincer Equation and the Second-Stage Results

Tables 4-A and 4-B show the results for the effect of NTR and ETR on wages, respectively. Columns (1) & (2) report the results from the Fixed-Effect estimation of equation (1) in which we directly include tariffs and union density in the Mincer wage equation. The First-Difference results from the Two-Stage estimation of equation (4) are presented in columns (3) & (4). We find a positive and significant interaction term between tariffs and union density under almost all scenarios. This finding is robust to using both NTR and ETR, and to controlling for trade outcome measures such as the lag of the import penetration and export intensity ratios. The use of lags alleviates endogeneity stemming from, arguably, the effect of labor costs on trade flows.

As previously mentioned, theories of political economy suggest the endogeneity of protection as product and labor market concerns are likely to be the basis upon which trade polices are formed. A country may be more inclined to maintain high levels of protection on industries with particular characteristics such as: higher absorbers of employment, high intensity of unskilled labor, with large union powers or with low levels of productivity. In Section 4.2 we discussed the decreasing cross industry dispersion in levels of protection as indicated by the declining standard deviations. This is confirmed in Figures 3-A and 3-b which plot the negative relation between the period changes in tariffs and pre-reform levels showing that proportionately larger tariff cuts were witnessed in industries with higher initial tariff level. This feature demonstrates the limited role of industry selectivity tariff reductions or lobbying during the reform period. Generally it is not easy to find a good instrument for tariffs, accordingly we address this concern in three different ways: (1) we employ lagged tariffs in our estimations, (2) we include industry fixed effects to control for the unobserved fixed industry characteristics affecting, simultaneously, wages and tariffs, and (3) we argue that 1993 pre-sample tariff rates are a good instrument for the yearly *changes* in contemporaneous tariffs. Regressing the change in NTR between 1993 and 2004 on the 1993 tariff rates yields a significant coefficient of -0.292 (t-value = 9.88) and an $R^2$ of 76.60% (similar results are obtained for ETR). Additionally, based on the South African government documents, it is believed that the acceleration in the tariff reduction program was in order to compensate for the depreciating Rand (Michie, 1997). Given that protection may respond to exchange rate pressures, we therefore also use the interaction of real effective exchange rate with 1993 tariffs as another instrument. This interaction with the yearly exchange rates allows our instrument to vary over time. Moreover given that gold dominated the pattern of trade in SA in the early nineties we also use the interaction between 1993 tariffs and gold prices

as another instrument. [17]

Results from using the latter IVs and their interactions with union density reported in columns (5) to (8) further confirm the positive and significant interaction of tariffs and union density. Columns (9) to (12) present results from the two estimation techniques using one period lagged tariffs instead of contemporaneous tariff rates. The latter scenario accounts for the possibility that wage adjustments may not occur instantaneously. Furthermore, as previously noted, it partially alleviates concerns with regards to tariff endoogeneity. Results using, both, ETR and ETR confirm our prior findings of a positive and significant interaction term. Another interesting result that is important to note is the negative coefficient on the tariff variable under some of the estimation scenarios. This finding implies that tariff cuts had a positive effect on wage premiums for industries with low union power. The significance of this effect is more robust to using ETR as our measure of protection.

To show the extent of our findings, we compare the effect of a one percentage point tariff decline on wages in the TRANS and TEX industries which had the highest period average rate of unionization of 55% & 47% respectively, to wages in the NoMet sector with the lowest unionization rate of 27%. Based on findings reported in column (6) of Tables 4-A, workers in the former two industries witnessed wage declines of 1.6% (-1.378+55%*5.372) & 1.14% (-1.378+47%*5.372) respectively due to a one percentage decline in NTR compared to a wage decline of 0.07% (-1.378+27%*5.372) for workers in latter industry. Considering the effect of cuts in ETR reported in Table 4-B, our results predict a wage reduction of 0.3% (-0.471+55%*1.509)& 0.23% (-0.471+47%*1.509) respectively for the TRANS & TEX industries versus a wage increase of 0.06% (-0.471+27%*1.509) for the NoMet industry. This impact of tariff cuts on wages is considerably large when compared to results found for Colombia where a one percentage point cut in NTR is estimated to reduce manufacturing wages by an average of 0.2% (Goldberg et al 2005). Yet these large effects are not surprising in light of the significantly high markups in SA as documented by Fedderke et al (2006) and Aghion et al (2006). The former work finds that markups in the South African manufacturing sector are approximately twice those reported for the US. Aghion et al (2006) show that markups are significantly higher in the South African manufacturing industries than they are in corresponding industries worldwide.

In Tables 5-A and 5-B we present regression results without controlling for collective bargaining. Differently, our estimates suggest that tariff cuts had a negative effect on wages as revealed by the positive and significant tariff coefficient estimate. This result is robust to using NTR

[17] A similar instrument was used in Goldberg et al (2004) where in the case of Colombia they use coffee prices

or ETR. To interpret our estimates in column (6), a one-percentage point reduction in NTR and ETR translate to 0.89% and 0.12% reduction in wage premium. These findings suggest the plausible omitted variable bias resulting from not controlling for industry heterogeneities in bargaining power.

So far we have discussed results from various estimation techniques: Fixed-Effect, First-Difference and IV, and for two specifications: controlling for tariffs and international prices and accounting for import penetration and export intensity. Our findings are also robust to other specifications in which we control for: (1) lagged values of exports and imports, (2) the latter interacted with real effective exchange rate to account for the differential the impact of exchange rates on industry premiums based on sectoral trade exposure, (3) sectoral capital intensity by controlling for the capital-labor ratio, (4) we also use SA's export price index as an alternative proxy for international prices, (5) our results are also robust to alternative means of aggregating our tariff variable. We reported findings from using pre-sample 1993 imports as weights to aggregate the tariffs rates. Our findings are also robust to using one period lag imports as weights or using simple averages[18], (6) finally, given that the potential endogeneity of trade policy may be in response to labor market concerns such as to protect low-skill intensive industries, we control for sectoral skill intensity using industry level employment data on workers' skill level from *Quantech*.[19]

# 7 Conclusion

Despite the number of empirical micro-data founded studies examining the effect of tariff reductions on industry wage premiums in developing countries, there has been no consensus on the nature of this relationship. Due to data limitations previous works fails to account for labor market institutions, such as union bargaining power, in spite of the theoretical foundation predicting a trade-wage-union relationship. Additionally, this literature only addresses reductions in NTR as the proxy for free trade policy. In this respect they do not control for the impact of changes in ETR that account for the *total* decline protection levels on both final output and the cost of raising protection on intermediate input.

In this paper we exploit the changes ETR and ETR in South Africa to investigate the role of collective bargaining in assessing the relationship between tariff changes and industry

[18]It is important to note that on the one hand import weighted tariffs give negligible weight to prohibitive tariffs given that the corresponding imports are typically low. On the other hand un-weighted average tariff rates may assign too heavy weights to commodities that are only a small fraction of imports.

[19]We do not report these results as they are very similar to those already documented.

wages. Using labor force data we cover the reform period from 1995 to 2004. Being privileged with a dataset that includes information on worker union membership status, we are able to control for industry union bargaining power in examining the wage-trade relationship. We find that the impact of tariffs is conditional on the industry's level of unionization. Only industries with higher union power were negatively affected by tariff cuts. This result is not surprising given that industries with stronger unions are able to accumulate larger shares of industry rents thereby securing their workers higher wages. With increased openness and intensified foreign competition industry rents are congested, consequently unions have less room to bargain for higher wages, a phenomena that explains the widespread protectionist sentiment across trade unions.

Policy implications based on our results, which reflect the South African trade reform experience, highlight the plausible role played by labor market institutions in the political economy of free trade. Our findings suggest that implementing free trade policy in countries with lower union power may have less of a social cost compared to more highly unionized nations. The fact that a country has a lower level of unionization may imply that workers are not capturing shares of economic profits, which in turn are accruing to firm owners. In such case it is probable that increased import competition would alternatively adversely affect the latter group. Noteworthy that in this paper the direct effect of tariffs on real wages takes place through the effect of free trade policy on consumer prices or on nominal wages. More efforts should be directed towards examining the indirect channels by which tariffs affect wages which can be through changes in industry productivity levels, producer prices or rents. Investigating the latter would compliment this work by fully understanding the dynamics of factor and product market imperfections in South Africa. Furthermore we believe that more country studies are needed to confirm the role of collective bargaining in the trade-wage relationship as predicted by this paper. Finally, an interesting avenue for future research would be to exploit the pool of existing empirical case studies for developing countries' trade reform episodes and theoretically modeling the witnessed diverging outcomes of free trade policies with respect worker wages. This challenging task would help identifying a set of pre-requisites that are needed to minimize the adverse effects of embarking on trade liberalizing policies.

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APPENDIX

Table 1

| Categories | Abbreviations |
|---|---|
| *Occupations:* | |
| 1) Legislators, senior officials and managers. | (Occupation 1) |
| 2) Professionals. | (Occupation 2) |
| 3) Technical and associate professionals. | (Occupation 3) |
| 4) Clerks. | (Occupation 4) |
| 5) Service and market sales workers. | (Occupation 5) |
| 6) Skilled agricultural and fishery workers. | (Occupation 6) |
| 7) Crafts and related trades workers. | (Occupation 7) |
| 8) Plant and machine operator and assemblers. | (Occupation 8) |
| 9) Elementary workers. | (Occupation 9) & (Omitted Category) |
| *Industries:* | |
| 1) Food, beverage and tobacco. | (FBT) |
| 2) Textile, clothing, leather products and footwear. | (TEX) |
| 3) Wood, paper, printing and publishing products. | (WPP) |
| 4) Coke, refined petrol, chemical, rubber and plastic products. | (PetChem) |
| 5) Glass and non-metallic minerals products. | (NoMet) |
| 6) Basic iron, steel, non ferrous, metal products, machinery and equipment. | (Met) |
| 7) Electrical machinery. | (EleMach) |
| 8) TV, radio and communication, scientific and professional equipment. | (RadTV) |
| 9) Motor vehicle, parts and accessories and transport products. | (TRANS) |
| 10) Furniture and other industries. | (FURN) |

Table 2
Tariff Data (%)

| | | Nominal Tariff Rates (NTR) | | | | | | Effective Tariff Rates (ETR) | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | ID | 1993 | 1997 | 2000 | 2004 | $\Delta$ (04-93) | $\widetilde{\Delta}$ (04-93) | 1993 | 1997 | 2000 | 2004 | $\Delta$ (04-93) | $\widetilde{\Delta}$ (04-93) |
| FBT | 1 | 25.97 | 14.07 | 14.60 | 11.70 | (14.27) | (11.33) | 96.03 | 51.076 | 45.37 | 38.42 | (57.61) | (29.38) |
| TEX | 2 | 52.18 | 36.98 | 29.14 | 19.78 | (32.40) | (21.29) | 187.78 | 131.89 | 81.08 | 53.43 | (134.35) | (46.68) |
| WPP | 3 | 15.28 | 7.32 | 6.51 | 6.24 | (9.04) | (7.84) | 30.01 | 13.455 | 9.16 | 9.49 | (20.52) | (15.78) |
| PetChem | 4 | 13.58 | 5.20 | 4.05 | 3.86 | (9.72) | (8.56) | 30.92 | 10.251 | 6.83 | 6.85 | (24.07) | (18.38) |
| NonMet | 5 | 17.42 | 6.59 | 6.06 | 6.15 | (11.27) | (9.60) | 38.47 | 13.163 | 11.21 | 11.70 | (26.76) | (19.33) |
| Met | 6 | 13.30 | 4.95 | 4.20 | 3.78 | (9.52) | (8.40) | 23.64 | 6.802 | 4.87 | 4.37 | (19.26) | (15.58) |
| EleMach | 7 | 21.16 | 7.45 | 7.69 | 7.15 | (14.01) | (11.56) | 46.02 | 14.095 | 14.74 | 13.84 | (32.18) | (22.04) |
| RadTV | 8 | 19.30 | 2.37 | 1.75 | 1.53 | (17.76) | (14.90) | 36.03 | -2.71 | -2.61 | -2.42 | (38.46) | (28.27) |
| TRANS | 9 | 24.62 | 15.93 | 13.65 | 11.62 | (13.00) | (10.43) | 61.59 | 40.522 | 24.82 | 21.78 | (39.81) | (24.64) |
| FURN | 10 | 27.94 | 9.50 | 7.75 | 7.07 | (20.86) | (16.31) | 50.44 | 15.99 | 19.89 | 18.46 | (31.98) | (21.26) |
| Wt.Avg. ✠ | | 19.62 | 9.35 | 7.93 | 6.68 | (12.94) | (10.82) | 47.75 | 22.72 | 15.36 | 12.79 | (34.96) | (23.66) |
| St.Dev. | | 11.44 | 10.00 | 7.99 | 5.27 | | | 49.46 | 39.354 | 24.74 | 16.80 | | |

$\Delta = \tau_{2004} - \tau_{1993}$
$\widetilde{\Delta} = \frac{\tau_{2004} - \tau_{1993}}{1+\tau_{1993}}$. Edwards(2005) argues this measure as more appropriate to capture the magnitude of changes in protection.
✠Tariffs are weighted by 1993 imports.

Table 3
First-Stage Estimation Results
Dependent Variable: Log Monthly Wages

| | 1995 | 1997 | 1998 | 1999 | 2000 | 2001 | 2002 | 2003 | 2004 | **Average ✠** |
|---|---|---|---|---|---|---|---|---|---|---|
| *Worker Characteristics:* | | | | | | | | | | |
| Age | 0.053 | 0.059 | 0.075 | 0.032 | 0.062 | 0.076 | 0.057 | 0.049 | 0.051 | 0.0575 |
| | (0.007)** | (0.009)** | (0.013)** | (0.012)** | (0.008)** | (0.009)** | (0.009)** | (0.010)** | (0.010)** | |
| Male | 0.256 | 0.272 | 0.357 | 0.236 | 0.340 | 0.256 | 0.263 | 0.318 | 0.269 | 0.2813 |
| | (0.026)** | (0.031)** | (0.052)** | (0.047)** | (0.031)** | (0.033)** | (0.034)** | (0.036)** | (0.034)** | |
| Household Head | 0.194 | 0.141 | 0.033 | 0.083 | 0.109 | 0.068 | 0.158 | 0.090 | 0.076 | 0.1162 |
| | (0.027)** | (0.032)** | (0.050) | (0.044)+ | (0.029)** | (0.031)* | (0.034)** | (0.035)* | (0.033)* | |
| Black | -0.729 | -0.695 | -0.811 | -0.512 | -0.876 | -0.830 | -0.742 | -0.809 | -0.699 | -0.7558 |
| | (0.035)** | (0.053)** | (0.090)** | (0.090)** | (0.053)** | (0.053)** | (0.057)** | (0.060)** | (0.066)** | |
| Other | -0.402 | -0.451 | -0.630 | -0.326 | -0.480 | -0.416 | -0.431 | -0.473 | -0.385 | -0.4327 |
| | (0.036)** | (0.058)** | (0.097)** | (0.099)** | (0.057)** | (0.059)** | (0.062)** | (0.066)** | (0.068)** | |
| Single | -0.083 | 0.008 | 0.160 | -0.094 | 0.099 | 0.092 | -0.080 | -0.083 | -0.115 | -0.0066 |
| | (0.028)** | (0.033) | (0.048)** | (0.045)* | (0.027)** | (0.029)** | (0.036)* | (0.035)* | (0.035)** | |
| *Worker Education:* | | | | | | | | | | |
| Not Complete Prim. School | 0.119 | 0.150 | 0.023 | 0.083 | 0.096 | 0.054 | 0.134 | 0.154 | 0.049 | 0.1011 |
| | (0.054)* | (0.060)* | (0.096) | (0.087) | (0.063) | (0.067) | (0.078)+ | (0.084)+ | (0.072) | |
| Complete Primary School | 0.307 | 0.356 | 0.204 | 0.289 | 0.325 | 0.258 | 0.353 | 0.297 | 0.199 | 0.2953 |
| | (0.045)** | (0.052)** | (0.082)* | (0.077)** | (0.055)** | (0.058)** | (0.068)** | (0.070)** | (0.062)** | |
| Complete Secondary School | 0.563 | 0.628 | 0.429 | 0.483 | 0.622 | 0.556 | 0.644 | 0.586 | 0.530 | 0.5730 |
| | (0.050)** | (0.059)** | (0.096)** | (0.088)** | (0.060)** | (0.064)** | (0.074)** | (0.075)** | (0.068)** | |
| University Degree or above | 1.015 | 1.012 | 0.137 | 1.424 | 0.978 | 1.114 | 1.038 | 1.025 | 1.153 | 1.0407 |
| | (0.107)** | (0.138)** | (0.344) | (0.206)** | (0.123)** | (0.125)** | (0.154)** | (0.161)** | (0.211)** | |
| *Worker Occupation:* | | | | | | | | | | |
| Occupation 1 | 0.750 | 0.475 | 0.455 | 0.628 | 0.782 | 0.811 | 0.769 | 0.788 | 0.922 | 0.6909 |
| | (0.060)** | (0.061)** | (0.106)** | (0.111)** | (0.081)** | (0.088)** | (0.085)** | (0.110)** | (0.109)** | |
| Occupation 2 | 0.614 | 0.433 | 0.584 | 0.591 | 0.643 | 0.485 | 0.874 | 0.494 | 0.724 | 0.5608 |
| | (0.146)** | (0.090)** | (0.245)* | (0.210)** | (0.125)** | (0.154)** | (0.191)** | (0.204)* | (0.227)** | |
| Occupation 3 | 0.640 | 0.397 | 0.522 | 0.325 | 0.468 | 0.539 | 0.586 | 0.557 | 0.488 | 0.5162 |
| | (0.053)** | (0.062)** | (0.116)** | (0.087)** | (0.056)** | (0.062)** | (0.064)** | (0.064)** | (0.070)** | |
| Occupation 4 | 0.395 | 0.289 | 0.323 | 0.250 | 0.419 | 0.375 | 0.488 | 0.455 | 0.368 | 0.3846 |
| | (0.042)** | (0.057)** | (0.089)** | (0.082)** | (0.053)** | (0.054)** | (0.061)** | (0.062)** | (0.066)** | |
| Occupation 5 | 0.067 | 0.142 | 0.004 | 0.036 | 0.129 | 0.028 | 0.219 | 0.066 | -0.205 | 0.0700 |
| | (0.067) | (0.068)* | (0.154) | (0.116) | (0.093) | (0.084) | (0.116)+ | (0.104) | (0.114)+ | |
| Occupation 6 | -0.175 | -0.096 | -0.524 | -0.745 | -0.059 | 0.071 | -0.180 | 0.452 | -0.119 | -0.1458 |
| | (0.204) | (0.100) | (0.330) | (0.261)** | (0.165) | (0.405) | (0.222) | (0.398) | (0.262) | |
| Occupation 7 | 0.227 | 0.106 | 0.080 | 0.067 | 0.031 | 0.110 | 0.124 | 0.137 | -0.042 | 0.1008 |
| | (0.033)** | (0.034)** | (0.057) | (0.053) | (0.036) | (0.039)** | (0.040)** | (0.042)** | (0.040) | |
| Occupation 8 | 0.253 | 0.194 | 0.220 | 0.119 | 0.217 | 0.194 | 0.197 | 0.230 | 0.215 | 0.2109 |
| | (0.028)** | (0.034)** | (0.057)** | (0.048)* | (0.033)** | (0.036)** | (0.038)** | (0.040)** | (0.038)** | |
| Observations | 2965 | 2304 | 1036 | 1232 | 2441 | 2073 | 1857 | 1641 | 1780 | 17329 |
| R-Squared (including ID & OD) | 0.560 | 0.410 | 0.460 | 0.380 | 0.500 | 0.530 | 0.520 | 0.490 | 0.450 | 0.478 |
| R-Squared | 0.450 | 0.290 | 0.380 | 0.270 | 0.370 | 0.380 | 0.370 | 0.340 | 0.280 | 0.348 |
| SD | 0.088 | 0.086 | 0.065 | 0.134 | 0.125 | 0.130 | 0.146 | 0.137 | 0.160 | |
| Year to Year Correlations | | 0.85 | 0.33 | -0.17 | 0.84 | 0.84 | 0.90 | 1.01 | 0.86 | |

Robust standard errors in parentheses, + significant at 10%; *significant at 5%; ** significant at 1%.
✠The average of the estimated parameters are weighted by inverse their respective variances.
*Other Included Controls:* age squared, 8 provincial dummies and 10 industry dummies.

Table 4-A
Nominal Tariff Rates - NTR

| Dependent Variable: | Log Monthly Wages | | Estimated Industry Premium | | | | | | Log Monthly Wages | | Est. Industry Prem. | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) | (11) | (12) |
| NTR | -0.412 | -0.468 | -0.363 | -0.550 | -1.498 | -1.378 | -0.807 | -0.689 | | | | |
| | (0.690) | -0.798 | (0.340) | (0.395) | (0.572)* | (0.694)+ | (0.615) | (0.720) | | | | |
| Lag NTR | | | | | | | | | -0.355 | -0.295 | -1.323 | -1.227 |
| | | | | | | | | | (0.667) | (0.665) | (0.368)** | (0.436)* |
| NTR*UD | 2.821 | 2.901 | 3.864 | 4.229 | 5.437 | 5.372 | 4.479 | 4.421 | | | | |
| | (1.113)* | (1.301)+ | (0.633)** | (0.694)** | (1.130)** | (1.225)** | (1.186)** | (1.240)** | | | | |
| Lag NTR*UD | | | | | | | | | 2.580 | 2.490 | 3.455 | 3.791 |
| | | | | | | | | | (1.061)* | (1.074)* | (1.097)* | (1.078)** |
| Union Density (UD) | 0.134 | 0.130 | -0.168 | -0.073 | -0.412 | -0.246 | -0.263 | -0.102 | 0.066 | 0.089 | -0.307 | -0.196 |
| | (0.252) | (0.274) | (0.189) | (0.202) | (0.313) | (0.302) | (0.266) | (0.262) | (0.272) | (0.269) | (0.290) | (0.313) |
| International Price | -0.002 | -0.002 | -0.002 | -0.003 | -0.003 | -0.003 | -0.002 | -0.003 | -0.002 | -0.001 | -0.003 | -0.004 |
| | (0.002) | (0.003) | (0.002) | (0.002) | (0.002) | (0.002) | (0.002) | (0.002) | (0.002) | (0.003) | (0.002) | (0.003) |
| Lag M Penetration | | -0.000 | | 0.000 | | 0.000 | | 0.000 | | -0.000 | | 0.000 |
| | | (0.000) | | (0.000)** | | (0.000)* | | (0.000)** | | (0.000) | | (0.000) |
| Lag X Intensity | | -0.041 | | -0.117 | | -0.129 | | -0.119 | | -0.055 | | -0.142 |
| | | (0.061) | | (0.052)+ | | (0.056)* | | (0.055)+ | | (0.065) | | (0.069)+ |
| Observations | 17328 | 17328 | 80 | 80 | 80 | 80 | 80 | 80 | 17328 | 17328 | 80 | 80 |
| R-squared | 0.52 | 0.52 | 0.32 | 0.38 | 0.29 | 0.36 | 0.31 | 0.38 | 0.52 | 0.52 | 0.19 | 0.26 |
| First Difference | N | N | Y | Y | Y | Y | Y | Y | N | N | Y | Y |
| Fixed Effect | Y | Y | N | N | N | N | N | N | Y | Y | N | N |
| Time Dummy | Y | Y | Y | Y | Y | Y | Y | Y | Y | Y | Y | Y |
| IV | N | N | N | N | Y | Y | Y | Y | N | N | N | N |

Robust standard errors in parentheses, + significant at 10%; *significant at 5%; ** significant at 1%.
We use the US export price index as the proxy for international prices.
Columns 1, 2, 9 &10 include individual controls: age, age squared, gender, single, household head, educational category, occupation and industry affiliation.
In columns 5 & 6 tariff changes are instrumented for by pre-sample tariffs, the real effective exchange rate interacted with pre-sample tariffs, and the latter interactions with union density.
In columns 7 & 8 tariff changes are instrumented for by pre-sample tariffs, gold prices interacted with pre-sample tariffs, and the latter interactions with union density.

Table 4-B
Effective Tariff Rates - ETR

| Dependent Variable: | Log Monthly Wages | | Estimated Industry Premium | | | | | | Log Monthly Wages | | Est. Industry Prem. | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) | (11) | (12) |
| ETR | -0.364 | -0.380 | -0.210 | -0.253 | -0.513 | -0.471 | -0.408 | -0.383 | | | | |
| | (0.147)* | (0.187)+ | (0.133) | (0.138)+ | (0.125)** | (0.140)** | (0.120)** | (0.131)* | | | | |
| Lag ETR | | | | | | | | | -0.358 | -0.342 | -0.237 | -0.234 |
| | | | | | | | | | (0.155)* | (0.153)+ | (0.133) | (0.151) |
| ETR*UD | 1.048 | 1.075 | 1.103 | 1.168 | 1.584 | 1.509 | 1.418 | 1.372 | | | | |
| | (0.264)** | (0.334)* | (0.207)** | (0.206)** | (0.251)** | (0.235)** | (0.255)** | (0.239)** | | | | |
| Lag ETR*UD | | | | | | | | | 0.983 | 0.956 | 0.779 | 0.853 |
| | | | | | | | | | (0.271)** | (0.273)** | (0.297)* | (0.288)* |
| Union Density (UD) | 0.057 | 0.054 | -0.087 | 0.016 | -0.289 | -0.121 | -0.219 | -0.066 | -0.006 | 0.016 | -0.104 | 0.020 |
| | (0.255) | (0.277) | (0.209) | (0.214) | (0.291) | (0.260) | (0.261) | (0.229) | (0.275) | (0.270) | (0.264) | (0.289) |
| International Price | -0.002 | -0.002 | -0.002 | -0.003 | -0.003 | -0.003 | -0.003 | -0.003 | -0.001 | -0.001 | -0.003 | -0.003 |
| | (0.003) | (0.003) | (0.002) | (0.002) | (0.002) | (0.002) | (0.002) | (0.002) | (0.003) | (0.003) | (0.002) | (0.003) |
| Lag M Penetration | | -0.000 | | 0.000 | | 0.000 | | 0.000 | | -0.000 | | 0.000 |
| | | (0.000) | | (0.000)** | | (0.000)* | | (0.000)** | | (0.000) | | (0.000) |
| Lag X Intensity | | -0.044 | | -0.100 | | -0.113 | | -0.108 | | -0.056 | | -0.139 |
| | | (0.068) | | (0.055) | | (0.056)+ | | (0.057)+ | | (0.068) | | (0.068)+ |
| Observations | 17328 | 17328 | 80 | 80 | 80 | 80 | 80 | 80 | 17328 | 17328 | 80 | 80 |
| R-squared | 0.51 | 0.52 | 0.29 | 0.34 | 0.26 | 0.32 | 0.28 | 0.33 | 0.51 | 0.51 | 0.18 | 0.25 |
| First Difference | N | N | Y | Y | Y | Y | Y | Y | N | N | Y | Y |
| Fixed Effect | Y | Y | N | N | N | N | N | N | Y | Y | N | N |
| Time Dummy | Y | Y | Y | Y | Y | Y | Y | Y | Y | Y | Y | Y |
| IV | N | N | N | N | Y | Y | Y | Y | N | N | N | N |

Robust standard errors in parentheses, + significant at 10%; *significant at 5%; ** significant at 1%.
We use the US export price index as the proxy for international prices.
Columns 1, 2, 9 &10 include individual controls: age, age squared, gender, single, household head, educational category, occupation and industry affiliation.
In columns 5 & 6 tariff changes are instrumented for by pre-sample tariffs, the real effective exchange rate interacted with pre-sample tariffs, and the latter interactions with union density.
In columns 7 & 8 tariff changes are instrumented for by pre-sample tariffs, gold prices interacted with pre-sample tariffs, and the latter interactions with union density.

Table 5-A
Nominal Tariff Rates - NTR

| Dependent Variable: | Log Monthly Wages | | Estimated Industry Premium | | | | | | Log Monthly Wages | | Est. Industry Prem. | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) | (11) | (12) |
| NTR | 1.065 | 1.057 | 1.407 | 1.384 | 0.887 | 0.898 | 0.815 | 0.814 | | | | |
| | (0.174)** | (0.158)** | (0.406)** | (0.404)** | (0.108)** | (0.118)** | (0.113)** | (0.105)** | | | | |
| Lag NTR | | | | | | | | | 1.062 | 1.065 | 0.240 | 0.345 |
| | | | | | | | | | (0.181)** | (0.153)** | (0.377) | (0.270) |
| International Price | -0.001 | -0.001 | -0.001 | -0.001 | -0.001 | -0.001 | -0.001 | -0.001 | -0.000 | -0.000 | -0.002 | -0.002 |
| | (0.003) | (0.003) | (0.001) | (0.001) | (0.001) | (0.002) | (0.001) | (0.002) | (0.003) | (0.003) | (0.002) | (0.002) |
| Lag M Penetration | | -0.000 | | 0.000 | | 0.000 | | 0.000 | | -0.000 | | 0.000 |
| | | (0.000) | | (0.000) | | (0.000) | | (0.000) | | (0.000) | | (0.000) |
| Lag X Intensity | | -0.027 | | -0.052 | | -0.060 | | -0.061 | | -0.053 | | -0.079 |
| | | (0.062) | | (0.059) | | (0.062) | | (0.062) | | (0.067) | | (0.064) |
| Observations | 17328 | 17328 | 80 | 80 | 80 | 80 | 80 | 80 | 17328 | 17328 | 80 | 80 |
| R-squared | 0.51 | 0.51 | 0.18 | 0.19 | 0.16 | 0.18 | 0.15 | 0.17 | 0.51 | 0.51 | 0.06 | 0.08 |

Table 5-B
Effective Tariff Rates - ETR

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) | (11) | (12) |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| ETR | 0.183 | 0.182 | 0.262 | 0.254 | 0.124 | 0.127 | 0.170 | 0.172 | | | | |
| | (0.036)** | (0.032)** | (0.057)** | (0.045)** | (0.043)* | (0.049)* | (0.028)** | (0.024)** | | | | |
| Lag ETR | | | | | | | | | 0.169 | 0.170 | 0.149 | 0.172 |
| | | | | | | | | | (0.040)** | (0.034)** | (0.072)+ | (0.063)* |
| International Price | -0.001 | -0.001 | -0.001 | -0.001 | -0.002 | -0.002 | -0.001 | -0.001 | 0.000 | 0.000 | -0.001 | -0.001 |
| | (0.003) | (0.003) | (0.001) | (0.002) | (0.002) | (0.002) | (0.001) | (0.002) | (0.003) | (0.003) | (0.002) | (0.002) |
| Lag M Penetration | | -0.000 | | 0.000 | | 0.000 | | 0.000 | | -0.000 | | 0.000 |
| | | (0.000) | | (0.000) | | (0.000) | | (0.000) | | (0.000) | | (0.000) |
| Lag X Intensity | | -0.029 | | -0.044 | | -0.060 | | -0.054 | | -0.061 | | -0.085 |
| | | (0.069) | | (0.064) | | (0.064) | | (0.066) | | (0.075) | | (0.062) |
| Observations | 17328 | 17328 | 80 | 80 | 80 | 80 | 80 | 80 | 17328 | 17328 | 80 | 80 |
| R-squared | 0.51 | 0.51 | 0.14 | 0.15 | 0.12 | 0.13 | 0.13 | 0.15 | 0.51 | 0.51 | 0.08 | 0.11 |
| First Difference | N | N | Y | Y | Y | Y | Y | Y | N | N | Y | Y |
| Fixed Effect | Y | Y | N | N | N | N | N | N | Y | Y | N | N |
| Time Dummy | Y | Y | Y | Y | Y | Y | Y | Y | Y | Y | Y | Y |
| IV | N | N | N | N | Y | Y | Y | Y | N | N | N | N |

Robust standard errors in parentheses, + significant at 10%; *significant at 5%; ** significant at 1%.
We use the US export price index as the proxy for international prices.
Columns 1, 2, 9 &10 include individual controls: age, age squared, gender, single, household head, educational category, occupation and industry affiliation.
In columns 5 & 6 tariff changes are instrumented for by pre-sample tariffs and the real effective exchange rate interacted with pre-sample tariffs.
In columns 7 & 8 tariff changes are instrumented for by pre-sample tariffs and gold prices interacted with pre-sample tariffs.

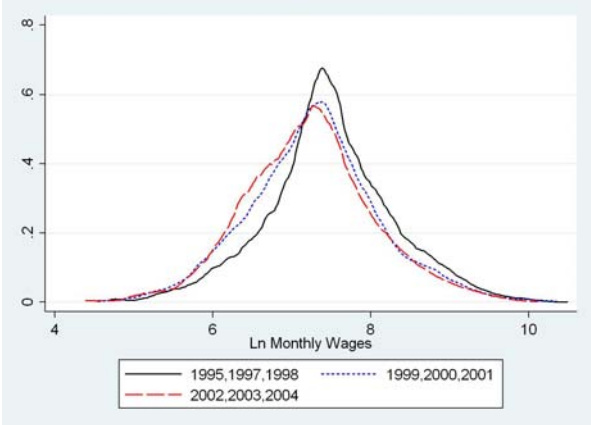


Figure 1: Wage Distribution (All Sample)

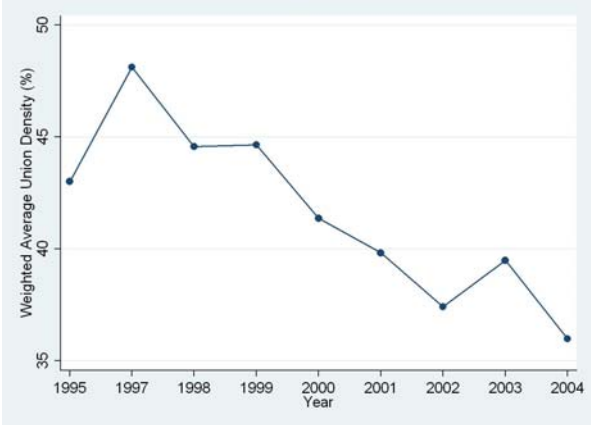


Figure 2: Employment Weighted Average Union Density (%)

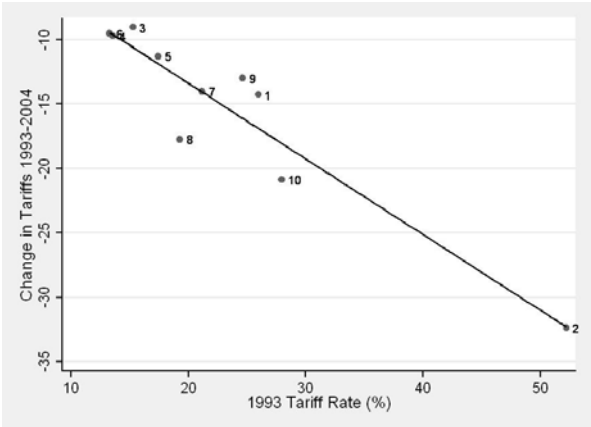


Figure 3-A: Changes in Nominal Tariff Rates

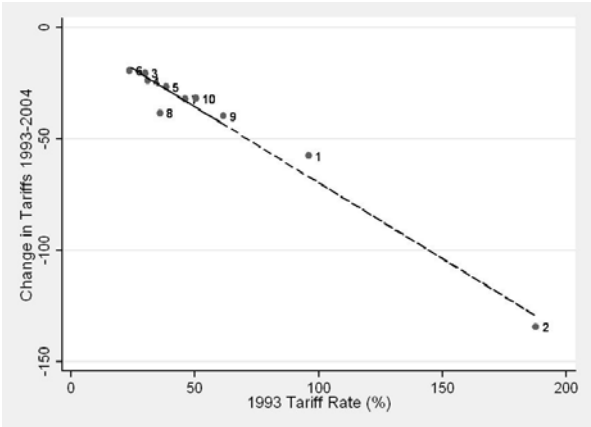


Figure 3-B: Changes in Effective Tariff Rates

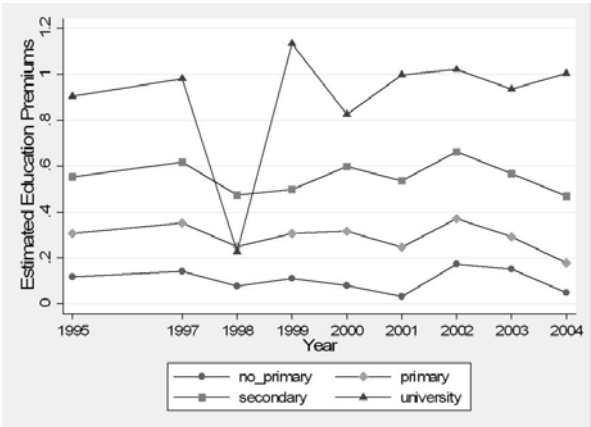


Figure 4: Estimated Education Premium

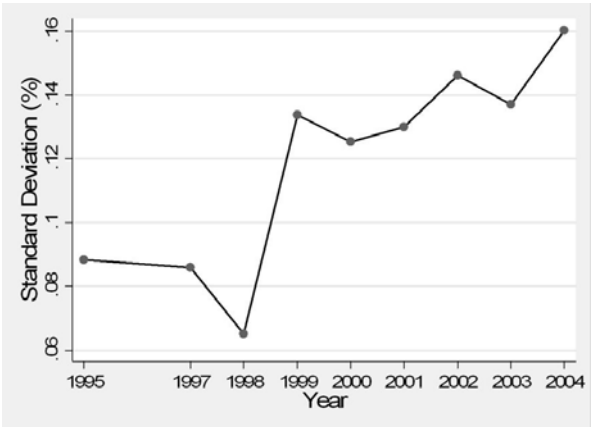


Figure 5: Weighted - Adjusted Standard Deviations of Estimated Industry Premiums