Department of Economics

Three Empirical Essays on Trade Reform in Post Apartheid South Africa

Riham Shendy

Thesis submitted for assessment with a view to obtaining the degree of Doctor of Economics of the European University Institute

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Dedication

To my family,
My father Mohamed Shendy, my mother Amal Zaatar, my brother Haytham, my sister-in-law Sherifa and my beautiful niece Jude. I hope this makes them proud.

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Part I

Introduction

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Introduction

This dissertation examines the trade reform experience in South Africa's (SA) manufacturing sector during the first post apartheid decade, from 1994 until 2004. During this period SA implemented an intensive trade reform policy particularly in the manufacturing sector. Average Nominal Tariff Rates and average Effective Tariff Rates 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. The core of this work explores the impact of this trade reform on the sector's labour and product market.

We note the significant amount of recent research focusing on SA. The importance of SA does not only stem from its special political setup. SA is an interesting case study in light of its particular status as a developing country. A common misrepresentation is categorizing the country as a developed economy. This misconception is attributed to its specific features that resemble developed economies, such as: GDP per capita levels of USD 11,110 (in PPP), higher than that reported for Latin American, North African and Sub Saharan countries that average to USD 7,014 , USD 4,500 and USD 2,792 respectively. Additionally the country has a thriving manufacturing sector that accounts for 20% of GDP, and an average rate of unionization of 35% which is comparable to levels is developed economies. Yet SA exhibits compelling developing country features. It ranks 121 out of 177 in the Human development indicator for 2006, lower than El Salvador, Nicaragua, Algeria, Honduras and Egypt. The Gini Index indicates higher levels of inequality than in Brazil, Mexico, Egypt and Argentina. The portion of South Africans living under the poverty line of USD 2 a day is 34%, much higher than in Brazil, Colombia, Mexico, Algeria and Morocco. Life expectancy is lower than in Bolivia, Egypt and Colombia, while health expenditure per capita is below that in Uruguay, Brazil and Lebanon. Furthermore, unemployment rates are estimated as high as 30%, while 50% of the workforce is defined as unskilled labour. In light of these characteristics it is natural to witness the attention dedicated by development economists to SA, particularly in light of the gold mine of macro and micro data available for long time periods, a scarce feature for developing countries.

In this dissertation We examine three novel questions regarding the impact of trade policy in SA. In Chapter 1 *Do Unions Matter? Trade Reform and Manufacturing Wages in South Africa* I investigate the effect of *nominal* tariff cuts on industry wage differentials using labor

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force data for the period from 1995 to 2004. This study extends 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.

In Chapter 2 *Efficiency Gains from Trade Reform: Foreign Input Technology or Import Competition? Evidence from South Africa* we extend the empirical trade literature examining the effect of trade reform, proxied by the reductions in Nominal Tariff Rates (NTR), on productivity, the latter estimated as the production function residual. We examine the different channels by which tariff cuts affect productivity growth. In the context of the South African trade reform experience, using industry level data for the manufacturing sector and covering the reform period from 1994 to 2004, we disentangle the differential effect of increased foreign competition, proxied by reductions in NTR, and that of the imported technology, proxied by the reductions in Input Tariff Rates (ITR), on productivity growth. Our measure of efficiency growth controls for the effect of tariff reductions on markups. The results suggest that the efficiency difference between foreign and domestic inputs have the major effect on productivity gains. Declines in ITR significantly raise productivity growth compared to an insignificant effect for NTR. Additionally, we find that higher protection rates are associated with higher markups, albeit this finding is not robust across all specifications.

In Chapter 3 *Heterogenous Trade Barrier Effects: What Can We Learn From a Disaggregated Gravity Model?* we relax the unrealistic assumption commonly imposed in gravity model estimations of equal elasticity of trade flows with respect to trade barriers across sectors and regions by using disaggregated industry level trade flows for South Africa's manufacturing sector and we test for heterogenous elasticities due to: (1) variations in the level of firm heterogeneity across sectors, and (2) differences in the size of trading partners. In line with theoretical predictions we find that the negative elasticity of *exports* to trade barriers increases for sectors with a higher level of firm homogeneity, supporting international trade models with heterogenous firms. Furthermore, we find that the negative elasticity of *imports* to trade barriers decreases for trading partners with larger export potential, emphasizing the objective of minimizing fixed costs of trading that is related to the number of countries with which South Africa, as an importer, trades.

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Part II

Chapters

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Chapter 1

Do Unions Matter? Trade Reform and Manufacturing Wages

1.1 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 the existing literature that investigates the effect of trade liberalization on wages premiums along two lines. Firstly, our work is the first to control for the effect of collective bargaining on wages when examining the trade-wage relationship. We hypothesize 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 measure captures the *total* impact of trade protection in an industry by accounting for both tariffs on final output and on intermediate inputs, accordingly ETR reflect the 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

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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 regard 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 in tariff cuts. This is demonstrated in the tariff data which show 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 time invariant 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

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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 cuts 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 significant reductions in wages. In more recent studies Goldberg et al (2005) and Pavcnik et al. (2004) examine the evolution of industry wage premiums in Colombia and Brazil, respectively, following the cuts in NTR. 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. Finally, using manufacturing census data for Indonesian firms, Amiti and Davis (2008) study the impact of reductions in both output and input tariffs on wages. The paper shows that cuts in output tariffs lower wages in import competing firms while boost wages in exporting firms. Moreover reductions in input tariffs increase wages in import-using firms relative to firms that source inputs locally.

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 reductions in 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

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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.

1.2 Country Background

1.2.1 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

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

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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]

1.2.2 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 35% in the period from 1995 to 2004. Of this percentage 70% were black & 64% male. This rate of unionization 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 at 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

[2] A thorough characterization of the tariff structure will be presented in Section 4.
[3] Prior to the new labor relations act of 1995 BC were known as Industrial Councils

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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 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 are 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 computed as the union density of an industry. With regard to the BC centralized effect on wages, we capture the latter by the industry, occupational and regional dummies.

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

1.3 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 that stems 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

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

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

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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.

1.3.1 The Correlation Between Trade Policy and Industry Union Power

Union Density, defined as the number of unionized employees in a specific industry over total employment of that industry, is the popular proxy for union bargaining power. Hirsch & Schumacher (2002) discuss two primary channels that affect this ratio. The first relates to the changes in levels of 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 trends 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 labor or female workers. For instance, an increase in skilled employment can be stimulated by trade induced Skill Biased Technological Changes, whereas rises in female participation can be due to labor reform which is 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 is 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

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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.*

1.3.2 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 wages of non-unionized workers 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 compression of industry rents, unions face a trade-off 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

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

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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.

1.4 Data

1.4.1 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 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 regard 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

[7]We find that the wage distribution and sample composition for 1994 is 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.

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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 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]

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

[9]Refer to Table 1 for abbreviations.

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1.4.2 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 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.

[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 = $Imports/(GDP + Imports - Exports)$.

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1.5 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 regard 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*.

1.5.1 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 (τ), 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.

$$Ln(w_{ijt}) = \beta_Q Q_{ijt} + \beta_\tau \tau_{jt} + \beta_{UD} UD_{jt} + \beta_{(\tau * UD)}(\tau_{jt} * UD_{jt}) + \beta_H H_{jt} + \beta_{ID} ID_{ijt} + \beta_{TD} TD_{ijt} + \epsilon_{ijt} \tag{1.1}$$

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 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.

1.5.2 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 policies suggesting the endogeneity of protection. As noted

[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.

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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 are a repeated cross section as opposed to a panel dataset of workers. Accordingly 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, β_{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 (1.3), 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 β_{ID} coefficients.[15] Given that our sample is composed of independent cross section data, we estimate equation (2) subject to (3) separately for each of the years in the sample.

$$Ln(w_{ijt}) = \beta_Q Q_{ijt} + \beta_{ID} ID_{ijt} + \epsilon_{ijt} \tag{1.2}$$

[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 regard 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.

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$$\sum_{j=1}^{10} n_{jt} * \beta_{ID,jt} = 0 \tag{1.3}$$

In the second-stage estimation we pool the yearly estimated industry wage premiums, β_{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 also control for time, fixed effects and P^* and estimate (4) in First-Difference:

$$\Delta\beta_{ID,jt} = \beta_\tau \Delta\tau_{jt} + \beta_{UD}\Delta UD_{jt} + \beta_{(\tau * UD)}\Delta(\tau_{jt} * UD_{jt}) + \beta_H \Delta H_{jt} + \nu_{jt} \tag{1.4}$$

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.

1.6 Estimation Results

1.6.1 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 regard to the education variables, the estimated coefficients are significant for most years and across most 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. 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

[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.

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:

$$SD(\beta) = \sqrt{\frac{\sum_j n_j(\beta_j - \beta_j^*)^2 - \sum_j n_j \sigma_{\beta_j}^2}{\sum_j n_j}} \tag{1.5}$$

where n_j is a vector of employment share in each industry, β_j and σ_{β_j} are the estimated industry premiums and their standard error, and β_j^* is the mean of β_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:

$$Correlation = \frac{\sum_j N_j[(\beta_{j,t} - \beta_{j,t}^*)(\beta_{j,t-1} - \beta_{j,t-1}^*) - \sigma_{\beta_{j,t}}\sigma_{\beta_{j,t-1}}]}{\sqrt{\sum_j N_j[(\beta_{j,t} - \beta_{j,t}^*)^2 - \sigma_{\beta_{j,t}}^2] * \sum_j N_j[(\beta_{j,t-1} - \beta_{j,t-1}^*)^2 - \sigma_{\beta_{j,t-1}}^2]}} \tag{1.6}$$

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 under study during which the reform took a smoother pace.

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1.6.2 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 policies are formed. A country may be more inclined to maintain high levels of protection on industries with particular characteristics such as: higher creators of employment, high intensity of unskilled labor, possess significant union bargaining power or have 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

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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 lag 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 regard 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 point 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 in 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.

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

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We note that our previous results are not driven by a plausible collinearity between the tariff and union density variables. Table 5-A and 5-B show the regression results of including both tariffs and union density and excluding the interaction term. The results support an independent and significant impact of each variable on wages. Moreover, computing the correlation matrix between tariffs (NTR and ETR) and union density reveals a correlation coefficient of -0.2 between the variables in *changes* and of 0.40 between the variables in *levels.* In Tables 6-A and 6-B we also present regression results without controlling for the role of union bargaining power and only including the tariffs. 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 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 *Quantec.*[19]

[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.

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1.7 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 limitation previous work fails to account for labor market institutions, such as union bargaining power, in spite of the sound 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 in protection levels by considering both tariffs on final output and on intermediate input in an industry.

In this paper we exploit the changes in NTR and ETR in South Africa to investigate the role of collective bargaining in assessing the relationship between tariff changes and industry 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 compressed, consequently unions have less room to bargain for higher wages, a phenomenon 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. We note 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. Other channels of transmission can be through changes in industry productivity, producer prices or rents. More efforts directed towards examining these channels is encouraged. Investigating

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the latter would complement 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 model 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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1.A Tables

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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	Δ (04-93)	$\widetilde{\Delta}$ (04-93)	1993	1997	2000	2004	Δ (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.

Shendy, Riham (2009). Three Empirical Essays on Trade Reform in Post Apartheid South Africa
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Table 3
First-Stage Estimation Results
Dependent Variable: Log Monthly Wages

	1995	1997	1998	1999	2000	2001	2002	2003	2004	**Avg ✠**
Worker Characteristics:										
Age	0.053	0.059	0.075	0.032	0.062	0.076	0.057	0.049	0.051	0.057
	(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	0.281
	(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	0.116
	(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	-0.755
	(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	-0.432
	(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	-0.006
	(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	0.101
	(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	0.295
	(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	0.573
	(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	1.040
	(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	0.690
	(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	0.560
	(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	0.516
	(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	0.384
	(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	0.070
	(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	-0.145
	(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	0.100
	(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	0.210
	(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 (inclu. 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.

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Table 4-A
Impact of 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	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)
Intern'al 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 Pen.		-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.

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Table 4-B
Impact of 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	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)
Inter'al 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 Pen.		-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.

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Table 5-A
Impact of 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	1.212	1.204	1.576	1.604	1.091	1.186	0.972	1.047				
	(0.145)**	(0.129)**	(0.373)**	(0.364)**	(0.168)**	(0.222)**	(0.175)**	(0.188)**				
Lag NTR									1.174	1.181	0.353	0.582
									(0.170)**	(0.149)**	(0.431)	(0.322)
Union Density	0.459	0.466	0.361	0.490	0.320	0.453	0.310	0.441	0.386	0.402	0.248	0.395
	(0.120)**	(0.120)**	(0.179)+	(0.185)*	(0.169)+	(0.169)*	(0.173)	(0.169)*	(0.136)*	(0.138)*	(0.201)	(0.197)+
Inter'al Price	-0.002	-0.002	-0.002	-0.003	-0.002	-0.003	-0.002	-0.003	-0.001	-0.001	-0.003	-0.003
	(0.002)	(0.003)	(0.001)	(0.001)+	(0.001)	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)	(0.002)	(0.002)
Lag M Pen.		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		-0.038		-0.086		-0.090		-0.092		-0.065		-0.113
		(0.058)		(0.055)		(0.059)		(0.061)		(0.064)		(0.069)
Observations	17328	17328	80	80	80	80	80	80	17328	17328	80	80
R-squared	0.52	0.52	0.22	0.26	0.21	0.25	0.20	0.25	0.51	0.52	0.08	0.13

Table 5-B
Impact of Effective Tariff Rates - ETR

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
ETR	0.204	0.203	0.289	0.286	0.163	0.185	0.197	0.215				
	(0.030)**	(0.027)**	(0.051)**	(0.036)**	(0.050)*	(0.064)*	(0.039)**	(0.040)**				
Lag ETR									0.185	0.187	0.156	0.194
									(0.037)**	(0.032)**	(0.071)+	(0.066)*
Union Density	0.368	0.373	0.317	0.429	0.279	0.401	0.289	0.409	0.313	0.335	0.243	0.389
	(0.133)*	(0.133)*	(0.169)+	(0.159)*	(0.169)	(0.160)*	(0.173)	(0.162)*	(0.156)+	(0.151)+	(0.181)	(0.187)+
Inter'al Price	-0.001	-0.001	-0.002	-0.002	-0.003	-0.003	-0.002	-0.003	0.000	0.000	-0.002	-0.003
	(0.003)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)	(0.002)	(0.002)
Lag M Pen.		-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		-0.038		-0.074		-0.084		-0.081		-0.073		-0.117
		(0.066)		(0.063)		(0.065)		(0.065)		(0.072)		(0.068)
Observations	17328	17328	80	80	80	80	80	80	17328	17328	80	80
R-squared	0.51	0.51	0.17	0.21	0.15	0.20	0.16	0.20	0.51	0.51	0.10	0.16
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.
In columns 7 & 8 tariff changes are instrumented for by pre-sample tariffs, gold prices interacted with pre-sample tariffs.

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Table 6-A
Impact of 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)
Inter'al 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 Pen.		-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 6-B
Impact of 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)*
Inter'al 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 Pen.		-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, the real effective exchange rate interacted with pre-sample tariffs.
In columns 7 & 8 tariff changes are instrumented for by pre-sample tariffs, gold prices interacted with pre-sample tariffs.

1.B Figures

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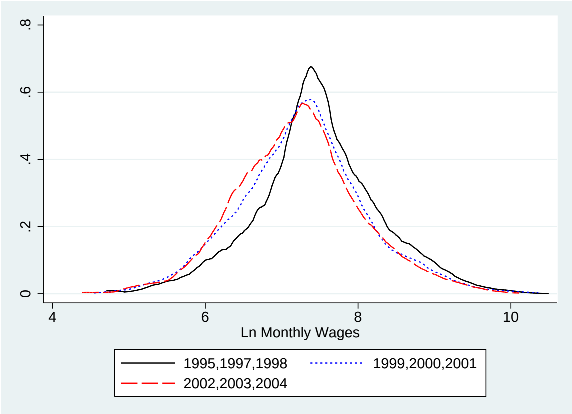

Figure 1: Wage Distribution (All Sample)

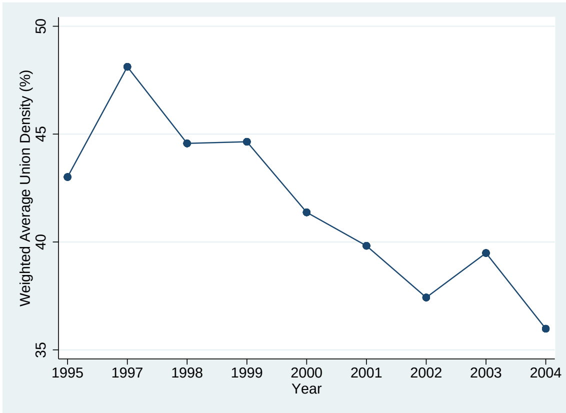

Figure 2: Employment Weighted Average Union Density (%)

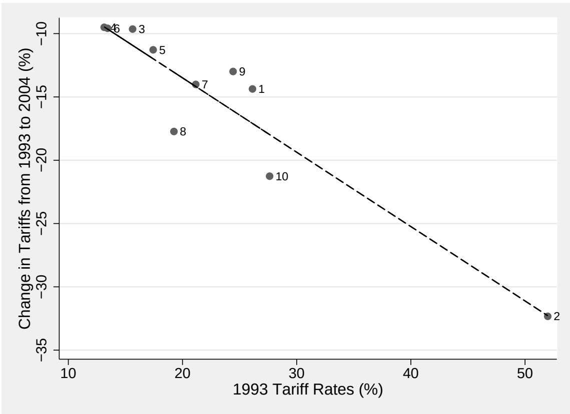

Figure 3-A: Changes in Nominal Tariff Rates in Percentage Points

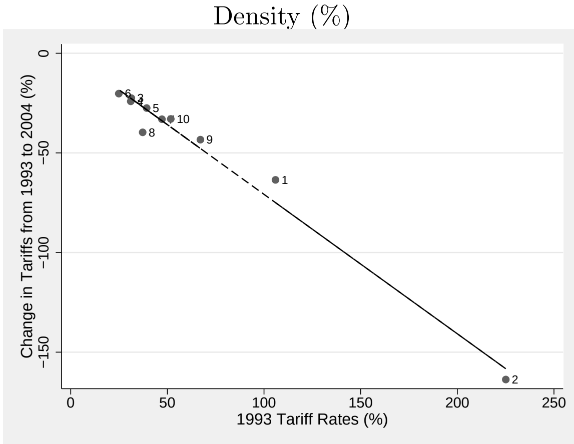

Figure 3-B: Changes in Effective Tariff Rates in Percentage Points

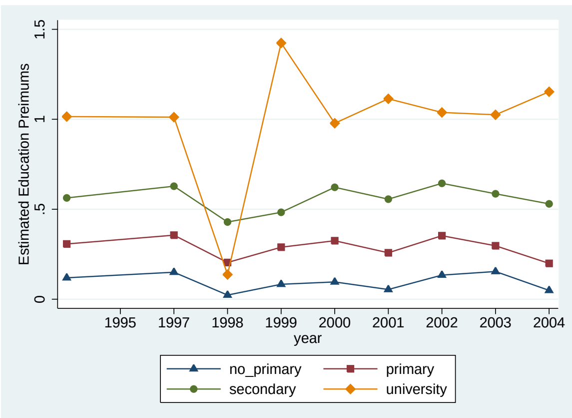

Figure 4: Estimated Education Premium

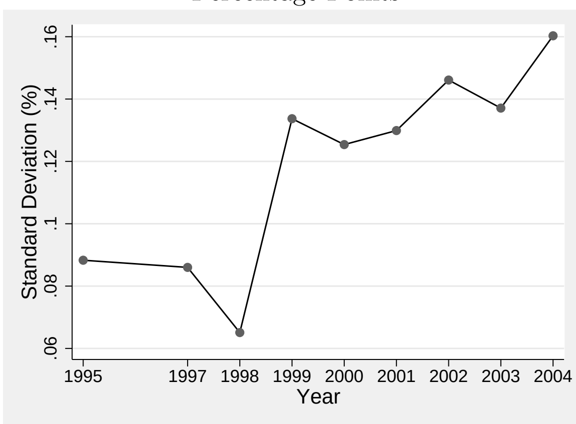

Figure 5: Weighted - Adjusted Standard Deviations of Estimated Industry Premiums

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Chapter 2

Efficiency Gains from Trade Reform: Foreign Input Technology or Import Competition?

2.1 Introduction

The South African (SA) manufacturing sector witnessed dramatic productivity growth between 1992 and 2000, coinciding with the increase in trade openness (Aron & Muellbauer 2007). In this paper we investigate the impact of tariff reductions on Total Factor Productivity (TFP) *growth* in SA using annual data for 28 manufacturing industries and covering the reform period from 1994 to 2004. We focus on two channels by which tariff cuts can induce productivity enhancement. The first occurs through the competitive pressures from the cheaper foreign imports of final goods, and the second takes place through the technological diffusion from the more accessible sophisticated imported inputs. We distinguish between these two channels by considering the differential effect of reductions in final output tariff rates, also known as Nominal Tariff Rates (NTR), and Input Tariff Rates (ITR). Reductions in NTR should capture the effect of increased competition in the domestic market, while reductions in ITR capture the effect of the new technology embodied in the cheaply available foreign inputs.

A common concern in the empirical literature investigating the effect of trade liberalization on "measured" productivity is the difficulty to distinguish between variations in productivity (or efficiency), from the markup squeeze often associated with exposure to free trade policies. In this paper we extend the existing literature by controlling for the latter, in addition to other concerns that might contaminate our inference on the effect of trade openness on efficiency growth. Firstly, we adopt two different econometric techniques to model industry markups

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and allow them to vary with changes in levels of protection. We then examine the differential effect of reductions in NTR and ITR on productivity growth. This procedure allows us to isolate productivity gains resulting for reduced price-cost margins from those induced by an actual increase in efficiency. The first technique we employ is based on Hall (1990) and entails a one step procedure in which we examine the effect of the tariffs on TFP growth and on markups simultaneously. The second procedure, based on Roeger (1995), involves a three stage estimation framework where we control for the variation in markups due to changes in tariffs in the first and second stages prior to estimating the effect of tariffs on efficiency growth in the third stage. The prime advantage of the latter technique lies in addressing the endogeneity concerns associated with the Hall approach. Secondly, we control for mis-measurements of primary inputs by including a proxy for capacity utilization. Additionally, we account for concerns with respect to the endogeneity of tariffs as suggested by theories of political economy. Generally it is not easy to find a good instrument for tariffs. Accordingly we address this concern by: (1) employing lagged tariffs in our estimations, (2) we include industry fixed effects to control for the unobserved time invariant industry characteristics affecting, simultaneously, productivity and tariffs, (3) we show that the structure of the SA tariff schedule during the reform period provides evidence that industry selectivity and lobbying were to some extent limited, (4) finally, we argue that SA's new government's liberal position in 1994 with respect to trade policy was triggered by their plan to reduce consumer prices and raise industrial efficiencies through curbing domestic monopoly powers who had vested interest in the prevalent protectionist policies (Bell 1997). Accordingly more reduced tariffs were applied to sectors with lower TFP. In light of our results where we find a negative coefficient on the tariff variable, fixing for this endogeneity bias will serve to further increase the magnitude of the negative impact of tariffs on the growth rate of productivity, hence our findings clearly imply that reductions in ITR induce increases in TFP growth.

Results from this work strongly suggest that it is through reductions in ITR that trade openness positively affects TFP growth. A one percentage point decline in ITR translates to 0.4% increase in productivity growth, compared to an insignificant effect of NTR. This finding implies that the efficiency difference between foreign and domestic inputs had a major effect on productivity gains in SA. We show that this result is robust across our two estimation procedures. Our findings are also robust to controlling for: the endogeneity of inputs, tariffs and the interaction of tariffs and markups; changes in capacity utilization over the business cycle; to using another measure of capital stock; and finally to employing Effective Tariff

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Rates (ETR) as an alternative proxy, to NTR, to capture the effect of increased import competition. With regard to the effect of tariff reductions on markups, our findings point to a decline in market power during the ten year reform period, but this result is not robust across all specifications. Finally, we note that a prime advantage of our data set is the availability of data on intermediate inputs. Controlling for the latter in our production function regressions ensures that the estimated effect of tariffs on productivity growth does not capture the increasing levels of imported materials due to the tariff reductions. Furthermore, accounting for intermediate input serves to control for the upward bias in markup estimates stemming from using value added data as opposed to gross output figures (Norrbin 1993, Hyde & Perloff 1995, Basu 1995).

The theoretical trade literature provides conflicting arguments regarding productivity gains attributed to trade reform in developing countries. Tybout & Westbrook (1995) outline the traditional arguments in support of such gains. Under imperfect competition, trade openness has *scale effects* as intensified foreign competition increases the price elasticity of demand, curbs domestic producers' market power, diminishes their markups, and ultimately increases "measured" productivity. Additionally increased competition under free trade policies can boost industry level productivity through the *share reallocation effect.* An industry wide increase in efficiency is witnessed if trade reform is associated with an increase in the market share of the more efficient firms and the exit of the less efficient ones. Finally, a *residual effect* can occur if trade policy positively affects firm productivity through unobserved channels such as innovation or technological progress.

On the other side of the debate, opponents of trade liberalization argue that conditions relevant to developing economies may prevent such gains to materialize. Pavcnik (2002) notes that gains from scale economies are not common in developing countries where increasing returns to scale are usually associated with import competing industries. With intensified foreign competition such industries are likely to contract. Rodrik (1988) shows that domestic firms are less likely to invest in catch-up technology that would reduce their costs (& increase their productivity) if trade liberalization decreases their domestic market share without increasing their international sales. Alternatively, foreign competition can negatively affect infant industries when learning by doing is important. Finally, the expected positive effect of increased competition on productivity due to efficient resource reallocation relies on the crucial assumption of free entry and exit of firms. Two market features prevalent in developing countries constitute serious obstacles to gains from such channel. On the one hand,

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the irreversibility of investment in capital equipment impedes the exit of less efficient firms. This concern is particularly relevant to countries which do not have well-developed secondary markets in capital equipment. On the other hand, binding credit constraints are likely to prevent the entry of new firms and the expansion of the existing efficient ones.

The empirical literature examining this issue has also failed to provide a consensus on the nature of the relationship between tariff reductions and productivity gains. Studies investigating the tariff effect on productivity *growth* produces conflicting results. Harrison (1994) finds an insignificant effect of tariff cuts on productivity growth in Cote d'Ivoire, while Tybout & Westbrook (1995) find that tariff reductions decreased productivity growth in Mexico by worsening scale efficiency. On the contrary Currie & Harrison (1997) find a significant positive effect of tariff reductions on productivity growth in Morocco. This effect is also confirmed by findings from Ferreira & Rossi (2003) and Muendler (2004) for Brazil. The evidence on the effect of tariff cuts on productivity *levels* seems more consistent as Pavcnik (2002), Topalova (2004) and Fernandes (2006) find that tariff reductions were associated with significant increases in TFP for Chile, India and Colombia, respectively.

While most of the aforementioned studies investigate the effect of tariff reductions on TFP (growth or level), the empirical methodology employed does not control for the plausible simultaneous change in markups during the reform period. Ignoring the latter is likely to produce biased estimates if one wishes to infer on the relationship between trade policy and efficiency. In this previous work, results suggesting an increase in TFP during the reform period can be due to reductions in markups as well as real efficiency gains. An exception to the aforementioned work is Harrison (1994) who explicitly models markups by using an econometric estimation that extends the Hall (1990) approach. Her findings suggest that trade openness lowered the price-cost margins, yet this effect is insignificant. Using the same procedure Levinsohn (1993) and Krishna & Mitra (1998) find that trade openness served to curtail markups in Turkey and India, respectively. One limitation to the aforementioned three studies is that they proxy trade openness by a time dummy that captures the reform period, in this respect the results do not account for the depth and the cross industry variations in trade policy. Moreover, using time dummies to account for the effect of changes in trade policy is likely to also capture the effect of other macro stabilization plans that took place during the same period. Furthermore, these studies employ balanced firm-level panel data in this respect results may suffer sample selectivity bias as the sample does not account for the plausible exit and entry of firms that might be triggered by the changes in trade policy.

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Employing a different methodology based on Roeger (1995) and using firm level data for European companies, Konings & Vandenbussche (2005) find that anti-dumping protection has a positive and significant effect on domestic markups.

Another common feature of the previous empirical work is using NTR as the proxy for changes in trade policy. The mechanism by which tariffs affect productivity has received little attention in this literature. As highlighted in Hallak & Levinsohn (2004), "a focus on mechanisms rather than just outcomes provides insight into choosing among the different flavors of trade policy to be able to evaluate when trade policy will be development policy". Accordingly, a more analytical investigation would entail distinguishing between NTR and ITR and examining their differential effects on productivity. On the one hand, reductions in NTR should reflect the competition effect of free trade policy on domestic production. Isolating the effect of increased foreign competition on markups, efficiency gains under this channel would result from the effect of competition on decreasing agency costs and eliminating managerial inefficiencies. Nickell et al (1997) summarize three channels through which competition reduces managerial slack. Firstly, a more competitive environment facilitates owners' ability to monitor managers due to the greater opportunities for comparison which can lead to sharper incentives. Secondly, increased competition raises the probability of facing bankruptcy which encourages managers to work harder to avoid such outcome. Thirdly, as competition raises demand elasticity, the reward to cost reductions increases, this enables lowering prices, increasing demand, and potentially higher profits. Alternatively, or simultaneously, import competition stimulated by the reduction of NTR may boost the overall productivity of an industry by forcing inefficient firms to exit the market. On the other hand, cuts in ITR increases efficiency by reducing the costs of foreign inputs. As noted in Tybout (2001), this enables domestic firms to expand their menu of intermediate inputs which allows each producer to match his input mix more precisely to the desired technology or product characteristic. Furthermore given the likely better quality-price ratio and advanced technological knowledge embodied in the imported input, productivity gains can be realized. Previous work investigating the hypothesis of the positive impact of foreign inputs on productivity focuses on testing the direct effect of using imported intermediate inputs on the latter as opposed to employing changes in ITR as a measure for the accessibility to foreign input. Muendler (2004) finds that foreign inputs played a minor role in productivity change in Brazil while Kasahara & Rodrige (2005) and Halpern et al (2005) find that imported inputs increased plant productivity in both Chile and Hungary, respectively.

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To our knowledge only two papers investigate the differential effect of NTR and ITR on "measured" productivity *levels*. Schor (2004) using data on Brazilian manufacturing firms finds that both increased competition and better access to imported inputs contribute to productivity gains in roughly the same magnitude. In another study Amiti & Konings (2007), using a census of manufacturing firms in Indonesia, find that the positive impact of reductions in ITR on productivity is three times the positive effect of reducing NTR. As noted in Schor (2004), changes in ITR serve as a better instrument to examine the impact of imported inputs on productivity due to two reasons. Firstly, imported inputs might be indirectly used by firms given that most manufacturing inputs usually undergo local remanufacturing. Secondly, the use of ITR enables directly testing the effect of a trade policy instrument as opposed to the impact of a trade policy outcome.

In this paper we also examine the differential impact of reductions in NTR & ITR on productivity growth in SA. As previously mentioned, the prime contribution of this paper lies in extending this emerging literature by controlling for a number of concerns. Firstly, we isolate the effect of tariff reductions on efficiency growth from that on markups. Secondly, we control for the mis-measurement in primary inputs by accounting for changes in capacity utilization. Thirdly, we address the issue of tariff endogeneity. And, finally, we control for intermediate inputs in our production function estimation.

We note that there is a recently growing literature employing firm-level data to investigate the effect of tariffs on productivity. Using disaggregated data are superior in controlling for firm heterogeneities within a sector. This may suggest some limitations associated with industry level analysis. We believe that sector-level analysis still serves to complement the aforementioned micro founded work in a number of aspects. Firstly, empirical work based on firm level data examines the effect of an industry level variable (tariffs) on a micro level outcome (firm productivity). Mountlon (1990) points to the bias from regressing a micro level variable on an aggregate variable due to the presence of intragroup error term correlations. Secondly, our industry level data are representative of SA's manufacturing sector and covers the whole reform period, in this respect it alleviates concerns regarding selectivity bias associated with firm level datasets. This is particularly relevant to studies investigating policy reform outcomes where it is important to account for sectoral expansions and contractions due to firm exit and entry. Thirdly, using industry-level real revenue deflated with the matching industry-level price deflators eliminates concerns associated with firm data that stem from the common procedure of deflating firm revenues with industry-level price deflators as op-

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posed to plant-level price data. This limitation to firm data analysis makes it impossible to differentiate between firm productivity differences and differences in markups. Finally, the noise attached to input and output firm level data constitutes a further concern. In the case that this noise is uncorrelated across units, aggregate measures are then perceived to be more precise.

This paper is divided into four sections. In the next section we present the empirical methodology. Section 3 describes SAs trade policy and the data. Section 4 discusses our results. And finally Section 5 concludes.

2.2 Methodology

In this section we outline the empirical methodology adopted in this paper to investigate the effect of tariff reductions on TFP growth controlling for the effect of the former on markups. Firstly we present the procedure commonly used in the empirical literature to investigate the relationship between tariffs and productivity and we briefly outline its limitations. Secondly, we model the relationship between tariffs and productivity growth using the Hall (1990) framework which allows a distinct modeling for the tariff effect on markups. We also present an extension to the latter methodology which allows an estimation procedure that does not require approximating an *unobserved* user cost of capital (Harrison 1994). In both settings we account for fluctuations in capacity utilization over the business cycle thus controlling for the bias attributed to the mis-measurement of inputs. Finally, given the plausible endogeneity bias associated with controlling for the effect of tariffs on markups under the previous techniques, we adopt the Roeger(1995) framework to address this concern.

2.2.1 The Common Approach

We start by assuming a production function for gross output in an industry (j) at time (t) of the following form:

$$Y_{jt} = A_{jt}F(K_{jt}, L_{jt}, M_{jt}) \tag{2.1}$$

Y, K,& M are the quantities of output, capital and intermediate input, respectively, proxied by their real values. L is labor employment and A is the industry specific index of Hicks-neutral technical progress or TFP. Taking *logs* of both sides [1] and then differentiating with

[1] Lower case letters indicate log variables, $y = log(Y)$.

respect to time we can re-express (1) as:

$$dy_{jt} = \varepsilon_k dk_{jt} + \varepsilon_l dl_{jt} + \varepsilon_m dm_{jt} + da_{jt} \tag{2.2}$$

where dy, dk, dm denote the annual growth rates in the real output, capital and intermediate input, dl is the growth rate of employment, ε_z is the elasticity of output Y with respect to an input Z, and da is the growth in TFP.

The commonly adopted procedure to estimate the impact of tariffs on productivity growth (or level) is to proceed in a two stage estimation framework. In the first stage a productivity estimate is derived and in the second stage the latter is regressed on tariffs. There are two alternative approaches to the first stage of the estimation strategy. The first is to compute productivity as the Solow residual which imposes the assumption of perfect competition and approximates output elasticities by factor shares. A prime shortcoming to this accounting decomposition is that it produces biased estimates in the presence of imperfect competition. The second approach is to treat equation (2) as a regression equation, estimate the three elasticity parameters and compute TFP as the regression residual. Despite the attractiveness of this procedure, and as highlighted in Basu & Fernald (2001), it involves the estimation of three parameters for output elasticities using data that often suffer multi-collinearity and are subject to differing degrees of endogeneity and thus OLS biases. Additionally the instrumental variable (IV) literature, which provides a partial solution to this problem, suggests the increasing problems of the IV approach in the presence of multiple endogenous variables. A crucial limitation to this estimation procedure is with regard to the estimated residual which captures both changes in markups and changes in efficiency. Accordingly a positive impact of tariff reductions on the residual can be attributed to either an increase in efficiency or merely a reduction in markups. Finally it is important to highlight that the robustness of the results in the second stage highly depends on the efficiency by which all three elasticity parameters are estimated in the first stage.

2.2.2 The Hall Technique

Hall (1990) provides a framework that allows one to address the above concerns. The main problem regarding the estimation of markups arises from the fact that while prices are measurable, marginal costs are unobserved. The Hall technique explicitly models markups by exploiting short-run fluctuations in output and production inputs. Employing this procedure

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enables one to empirically disentangle the effect of trade policy on efficiency from that on market power. Additionally, the Hall procedure is preferable to a regression estimation of (2) as it requires estimating only one parameter of the production function, the markup, as opposed to three input parameters. This feature increases efficiency and produces better IV estimates.[2]

Under monopolistic competition firms charge a price that is a markup, μ, over marginal cost. The first order condition for profit-maximization under imperfect competition implies that $\frac{dY}{dZ} = \mu \frac{P_Z}{P_Y}$ for an input Z. Multiplying both sides by the input-output ratio $\frac{Z}{Y}$ we can re-express the elasticities in (2) as a markup multiplied by each input's share in gross output:

$$dy_{jt} = \mu[\alpha_k dk_{jt} + \alpha_l dl_{jt} + \alpha_m dm_{jt}] + da_{jt} \tag{2.3}$$

where α_l & α_m are the shares of nominal labor remuneration and intermediate inputs, respectively, in nominal gross output, computed from the data. To estimate the markup parameter in equation (3) we face three concerns. The first concern is with regards to calculating the capital factor share $\alpha_k = \frac{rP_K K}{P_Y Y}$ where $P_K K$ & $P_Y Y$ are nominal capital stock and gross output respectively. To compute the latter ratio one needs to estimate an *unobserved* user cost of capital, r. In line with other work (Aghion et al 2006, Fedderke et al 2005, Ferreira & Rossi 2003, Griffith et al 2005, Oliveira et al 1996) we approximate r by the long run nominal interest rate less expected inflation plus depreciation.[3] A second problem in estimating equation (3) concerns the computational choice of the factor shares. One option would be to assume shares that are constant over time and compute average shares which represent steady state values. Alternatively, one can allow factor shares to vary period-by-period and adopt the Tornquist approximation, $\frac{\alpha_t + \alpha_{t-1}}{2}$, which relaxes the assumption of constant output elasticities. Basu & Fernald (2001) thoroughly argue the pros and cons of each technique. Following their work we also use the former approximation. The third concern which is common in production function estimations is the endogeneity of inputs. Inputs and output are simultaneously determined by the firm, accordingly the technical change term, da, is correlated with the choice of inputs. Ignoring the latter leads to biased OLS estimates. The

[2] As noted in Basu & Fernald (2001) and as will be discussed, this advantage comes at a cost of imposing a profit maximization assumption. This is a relatively weak condition that one expects to approximately hold.

[3] For the interest rate we use ten year government bond yields. Expected inflation is based on the CPI and is computed using the Hodrick-Prescott filter. The depreciation rate is set to 10% which is equivalent to an average service life of 10 years. Our results are also robust to using a rate of 5%. Due to the lack of data we are unable to construct a sector specific user cost of capital similar to that used in Hall (1990).

common approach to overcome this problem is to use IVs that are correlated with inputs but independent from any demand or productivity shocks that affect the firm. Following Arellano & Bond (1991) we use lagged values of k, l & m as IVs. In light of the above framework we allow tariffs to affect both markups and TFP and proceed to our final estimating equation:

$$dy_{jt} = \mu[dx_{jt}] + \mu_{ETR}[ETR_{jt} * dx_{jt}] + \gamma_{NTR}NTR_{jt} + \gamma_{ITR}ITR_{jt} + \gamma_j + \gamma_t + d\eta_{jt} \qquad (2.4)$$

where dx refers to the term in the bracket in equation (3) capturing the growth in input weighted by their respective shares in gross output. If the manufacturing sector in SA exerts market power then we expect our estimate of μ to exceed one. To allow markups to vary with protection and given that we are concerned with the effect of changes in *total* protection on markups, we control for an interaction between dx and Effective Tariff Rates (ETR). The latter measure of tariffs combines the net effect of both NTR and ITR on an industry; in this respect it measures the *total* effect of protection, defined as:

$$ETR_j = \frac{\tau_j - \sum_{i=1}^{N} b_{ij}\tau_i}{1 - \sum_{i=1}^{N} b_{ij}} \qquad (2.5)$$

where τ_j is NTR in an industry j, τ_i is NTR in an industry i, b_{ij} is the free trade technical coefficient constructed from the input-output table and measures the share of an input i in the cost of an output j at free trade prices[4], and $b_{ij}\tau_i$ is the ITR facing an industry j. We expect μ_{ETR} to be positive if higher protection is associated with higher industry market power. The coefficients on the NTR and ITR in equation (4) capture the effect of increased import competition and the impact of increased access to foreign inputs, respectively, on efficiency growth. In our estimation we use one period lags of NTR, ITR and ETR. In line with Fernandes (2006) & Topalova (2004) we estimate the effect of lagged tariffs rather than contemporaneous values to account for the possibility that productivity adjustments may not occur instantaneously. Furthermore the latter specification partially alleviates concerns regarding the endogeneity of protection. If productivity growth increases due to reductions in either NTR and ITR, then γ_{NTR} and γ_{ITR} should be negative. Finally, we control for industry fixed effects, γ_j, to allow for sector specific markups and technologies that are constant over time. Industry dummies also serve to control for the plausible endogeneity of tariffs as they capture the unobserved time invariant industry characteristics affecting, simultaneously, productivity and tariffs. We

[4]The unobserved free trade coefficient is defined as: $b_{ij} = \hat{b}_{ij} * \frac{1+\tau_j}{1+\tau_i}$, where $\hat{b}_{ij}$ is the observed value of an input i in the gross value of an output j, under protection.

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also include time fixed effects, γ_t, to capture period related macroeconomic factors such as privatization or any other stabilization plans.

The Harrison Extension

To check the robustness of our results and to ensure that our findings are not contaminated by a plausible mis-approximation of the user cost of capital, r, we employ the methodology of Harrison (1994) which extends the Hall procedure. The technique exploits the feature that the scale elasticity parameter, β, is the sum of the elasticities of output with respect to inputs. Using the profit maximization first order condition, $\varepsilon_z = \mu\alpha_z$, we can substitute α_K in equation (3) by $(\frac{\beta}{\mu} - \alpha_l - \alpha_m)$. With some algebraic manipulation we obtain estimation equation (6) which entails estimating a markup, μ, and a scale parameter, β:

$$dy_{jt} = \mu[\alpha_l d\widetilde{l_{jt}} + \alpha_m d\widetilde{m_{jt}}] + \beta dk_{jt} + da_{jt} \tag{2.6}$$

$\widetilde{l}$ & $\widetilde{m}$ are equal to $ln(\frac{L}{K})$ & $ln(\frac{M}{K})$ respectively, and dy & dk are the growth rates in real output and capital stock. Allowing for trade policy to affect efficiency growth and markups, similar to equation (4), our final estimating equation under this scenario is:

$$dy_{j,t} = \mu[d\widetilde{x_{jt}}] + \mu_{ETR}[ETR_{jt} * d\widetilde{x_{jt}}] + \gamma_{NTR}NTR_{jt} + \gamma_{ITR}ITR_{jt} + \beta dk_{jt} + \gamma_j + \gamma_t + d\eta_{jt} \tag{2.7}$$

Controlling for Capacity Utilization

The models presented in the previous sections are based on the assumption that we correctly observe capital services K and labor input L. An additional source of bias in the assessment of the tariff-productivity relationship comes from the plausible mis-measurement of inputs. In practice labor and capital input may fluctuate as capacity utilization changes over the business cycle. Under such condition the observed *number* of workers and *quantity* of capital do not reflect the intensity of factor use. To account for the latter, under the Hall estimation we adopt the technique outlined in Basu & Fernald (2001), originally attributed to Flux (1913). We break intermediate input into two components, the flow of energy input (E) and all the other intermediate input (O), and we extend (3) to:

$$dy_{jt} = \mu[(\alpha_k + \alpha_e)de_{jt} + \alpha_l dl_{jt} + \alpha_o do_{jt}] + da_{jt} \tag{2.8}$$

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where e is $log(E)$; log of the real value of inputs of Electricity, Gas and Water. To control for capacity utilization in the Harrison framework and similar to Harrison (1994) we model capital services as equal to K*E. Accordingly in equation (6) $\widetilde{l}$ & $\widetilde{m}$ are redefined as equal to $ln(\frac{L}{K*E})$ & $ln(\frac{M}{K*E})$ respectively, and dy & dk are the growth rates in Y & $(K * E)$, respectively.[5]

2.2.3 The Roeger Technique

In Section 2.2 we presented a one-step estimation procedure in which we examine the effect of trade openness on productivity growth and on mark-ups simultaneously. We address the issue of the endogeneity of inputs by considering lagged values of k, l & m as IVs for dx. With regard to equations (4) & (7), it is not clear in the IV literature how to instrument for the interaction of dx and ETR. This implies that our previous results might suffer endogeneity bias. In this section we outline an alternative three stage procedure that accounts for the latter concern. In principle this procedure is similar to the traditional two-stage estimation technique in Section 2.1 in which the production function residual from the first stage is regressed on the tariff variable in the second stage; yet differently this technique allows controlling for the variation in markups due to changes in ETR in the first and second stages prior to estimating the effect of tariffs on efficiency growth (the residual) in the third stage.

In the first stage we employ the Roeger(1995) methodology which estimates markups while overcoming the identification problems arising from the correlation between inputs and the error term. Under the assumptions of constant returns to scale (CRS), perfect competition, and the absence of labor hoarding or capital under-utilization, both the residuals from the production function (Solow Residual-SR) in equation (9), and the price-based residual from the cost function (Dual Solow Residual-DSR) in equation (10) are highly correlated, where B is the Lerner index $(1 - \frac{1}{\mu})$ and p, q & r are log prices of Y, L & K respectively. Under the assumption of constant returns to scale, Roeger (1995) shows that a lack of correlation between the SR & the DSR is a consequence of positive markups rather than the presence of fixed factors of production.

[5] Unfortunately we do not have data on labor input in hours thus we can not control for the variation of labor intensity across the business cycle. We hypothesize that our energy proxy will also capture the latter particularly given our emphasis on the manufacturing sector.

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$$SR = dy_{jt} - \alpha_l dl_{jt} - \alpha_m dm_{jt} - (1-\alpha_l-\alpha_m)dk_{jt} = B(dy_{jt} - dk_{jt}) + (1-B)da_{jt} \quad (2.9)$$
$$DSR = -[dp_{jt} - \alpha_l dw_{jt} - \alpha_m dq_{jt} - (1-\alpha_l-\alpha_m)dr_{jt}] = -B(dp_{jt} - dr_{jt}) + (1-B)da_{jt} \quad (2.10)$$

By subtracting (10) from (9) the unobserved productivity shocks, da, cancels out and we obtain equation (11):

$$(dy+dp)_{jt} - \alpha_l(dl+dw)_{jt} - \alpha_m(dm+dq)_{jt} - (1-\alpha_l-\alpha_m)(dk+dr)_{jt} = B[(dy+dp)_{jt} - (dk+dr)_{jt}] \quad (2.11)$$

Rearranging (11) we obtain (12) and we can directly estimate the markup, μ, by simple OLS:

$$dz_{jt} = \mu[dh_{jt}] \quad (2.12)$$

where:

$$dz_{jt} = d(y+p)_{jt} - d(k+r)_{jt} \quad (2.13)$$
$$dh_{jt} = [\alpha_l d(l+w)_{jt} + \alpha_m d(m+q)_{jt} - (\alpha_l+\alpha_m)d(k+r)_{jt}] \quad (2.14)$$

$d(y+p)$, $d(l+w)$ & d(m+q) are the log change in nominal gross output, labor costs and intermediate input respectively, $d(k+r)$ is the log change in the users cost of capital multiplied by nominal capital. Note that the Roeger technique is based on the assumption of constant returns to scale. Oliveira et al (1996) show that if the latter assumption is dropped, the estimated markup, $\widehat{\mu}$, captures the ratio of markup to the scale elasticity parameter ($\widehat{\mu} = \frac{\mu}{\beta}$). Thus in the case of increasing returns to scale $\widehat{\mu}$ should be interpreted as the lower bound value of the markup. In Section 4 we will show that results from using the Harrison procedure yield an estimate of the scale parameter, $\widehat{\beta}$, that is not statistically different from one, indicating the presence of constant returns to scale. This implies that the coefficient $\widehat{\mu}$ estimated under this methodology is most likely an unbiased estimate of the markup. In the context of our work we allow the markup to vary with changes in lagged ETR and we estimate equation (15) by OLS:

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$$dz_{jt} = \mu[dh_{jt}] + \mu_{ETR}[dh_{jt} * ETR_{jt}] \tag{2.15}$$

$$dy_{jt} = \mu[dx_{jt}] + \mu_{ETR}[ETR_{jt} * dx_{jt}] + d\nu_{jt} \tag{2.16}$$

In the second stage we compute an adjusted Solow Residual. We substitute our estimates of μ & μ_{ETR} obtained from (15) in equation (16) and we simply *compute* efficiency growth as the residual $d\nu_{jt}$. Different from the residual da_{jt} from regression equation (3), dv_{jt} captures productivity growth net of the effect of total protection, proxied by ETR, on markups. In the third stage we regress the computed $d\nu_{jt}$ on the lag of NTR & ITR, time dummies and industry dummies. [6]

2.3 South Africa: Trade Policy and the Data

Up until the 1970s SA was firmly oriented towards import substitution industrialization. The latter consisted of a wide-ranging system of quantitative restrictions as opposed to tariff-based protection. The first shift away from this trade regime came in 1972 with the relaxation of quantitative restrictions and the introduction of an Export Development Assistance scheme; however the overall trade policy remained protectionist. Starting in 1985 and as quotas were replaced by equivalent import tariffs, SA faced balance of payment pressures arising from the debt crisis and from capital outflows due to foreign disinvestment and sanctions. In an attempt to maintain current account surpluses in excess of the required foreign debt, SA's government imposed import surcharges. The latter led to an increase in tariff rates thus offsetting the effects of the relaxations of quotas. Belli et al. (1993) find that by the end of the 1980s the coefficient of variation of SAs tariffs was the second highest of 32 developing countries. In April 1994, the first post-apartheid government was democratically elected. This coincided with the initiation of multilateral trade reform as the WTO agreed on the phase-down tariff plan offered by SA in the GATT/WTO Uruguay Round. By signing the latter agreement the country committed itself to the rationalization of tariff lines, removal of quotas and export subsidies. Consequently, starting in 1995 SA experienced considerable cuts in protection rates.[7]

[6] Note that we do not control for changes in capacity utilization under this procedure. We believe that this does not affect our results given that controlling for the latter under both the Hall and Harrison procedures does not significantly alter our findings.

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

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In this paper we measure changes in trade policy using annual applied NTR assembled by Edwards (2005a). The data covers 28 manufacturing industries at the SIC-3 digit level and covers the period from 1988 to 2004. We consider the data from 1994 to 2004. Narrowing down the period under study serves our purposes of investigating SA's performance during the trade reform period which started in 1994. Furthermore it restricts our analysis to the particular political time frame of the first post apartheid regime. Additionally, confining our analysis to using data from the mid nineties provides a more consistent tariff series as Edwards (2005a) is unable to estimate the *ad valorem* equivalent of the Non Tariff Barriers which were still prevalent in some sectors prior to 1994. This suggests that the computed tariffs prior to the mid 1990s may be underestimating protection.

While NTR measure protection on final output, Edwards (2005a) also estimates ETR which measures the *total* effect of protection on output by accounting for tariffs on final output in addition to those imposed on intermediate inputs in an industry. As noted in Section 2, ITR in an industry j are constructed as a weighted average of NTR on inputs, i, that enter in the production process of j. The weights are based on cost shares for 42 input industries.[8] For example, if agricultural input accounts for 35% of the food industry gross output, while other inputs from the food industry and the services sector contribute to 12% and 15% respectively of the food industry gross output, then the ITR on food is equal to 35% of the NTR on agriculture plus 12% of the NTR on food plus 15% of the NTR on services, the latter assumed to be zero since it is a non traded input.[9] We note that the level of industry aggregation used to construct the ITR implies a potential bias in the estimated effect of NTR on productivity as part of the NTR effect will operate through our measure of ITR. This will be discussed in more detail in the next section.

Table 1 summarizes our tariff data. Average NTR & ITR decreased from 26% & 9% respectively in the pre-reform year 1993 to 9% & 3.5% in 2004. The tariff cuts were not uniform across sectors where industries such as wearing apparel, tobacco and footwear experienced the highest reductions in NTR, and industries such as wearing apparel, textile, communication equipment and footwear witnessed the highest cuts in ITR. Notably tariff rationalization was

[8] As noted in Edwards (2005), the ITR for an industry j is constructed as the weighted average of NTR on all inputs i used to produce j. The weights are constructed using Input-Output (IO) and Supply-Use (SU) tables which are reduced to 42 industrial sectors based on the SIC classification system. ITR for 1988 and 1989 are drawn from the IO tables and ITR for 1993, 1998, 1999 and 2000 are drawn from the SU tables. The interim years are calculated as a weighted average using the ITR of the two tables that bound the period (with a linearly declining weight).

[9] Giving non-traded products a zero tariff rate is in line with Balassa (1965).

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intense in the start of the reform period, leveling out from 1999 as average rates reached less than 50% of their 1993 pre-reform value (Figures 1). This change in trade policy translated into significant increases in trade flows. In Figures 2 & 3 we plot the trade openness ratio and real imports respectively. The figures demonstrate the significant increase in trade volumes during the reform period compared to the earlier period. In the bottom row of Table 1 we see the decline in the yearly standard deviations of both NTR and ITR indicating the decreasing cross industry dispersion of protection rates. The latter is also displayed in Figure 4 which plots the pre-reform NTR in 1993 against the change in NTR between 1993 and 2004 as computed in column (5) of Table (1). The downward sloping graph reveals that industries with initially high levels of protection experienced more severe liberalization measures.[10] This feature of the data implies that industry selectivity reduction or lobbying were to some extent limited as all tariffs reached similarly low levels.

To estimate productivity we use a panel dataset on gross output, intermediate input, capital and employment from the South African Standardized Industry Database (Quantec Research, 2006). The data are provided at the 3-digit SIC classification for the period from 1970 to 2005. The data are available at both current and 2000 constant prices. A prime advantage to this data set lies in the ability to control for intermediate inputs in the production function regression which ensures that the estimated effect of tariffs on productivity does not capture the increasing levels of imported materials due to the tariff reductions. Additionally, accounting for intermediate input serves to control for the upward bias in markup estimates stemming from the use of value added figures as opposed to gross output data.

2.4 Results

Table 2 shows results from the three estimation procedures outlined in Section 2. As previously noted we estimate the effect of one period lagged tariffs in order to take into account that productivity adjustments may not occur instantaneously. Moreover using lags partially alleviates concerns with regard to the endogeneity of protection. Columns (1) & (2) refer to the results that arise from using the Roeger procedure in which we control for both the endogeneity of inputs and the endogeneity of the interaction between inputs and ETR. Column (1) presents the first stage estimation results from equation (15) from which we estimate a markup and the impact of ETR on the latter. In column (2) we show the Roeger third stage

[10] Similar graphs are obtained for ITR and ETR.

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estimation findings from regressing the computed efficiency growth (net of the effect of tariffs on the markup), from equation (16), on the tariff variables.

The point estimate of the markup is 1.144 (see columns (1)). The presence of rents in the South African manufacturing sector is confirmed by our test on the coefficient as we reject the null of $\mu = 1$ at a 5% significance level (as shown in the bottom row of Table 2). This feature of market power in SA is well documented in a number of works (Fedderke et al 2006, Aghion et al 2006, Edwards et al 2005b). With regard to the relationship between *total* protection and markups, as captured by the interaction between ETR_{jt} and dx_{jt}, our results support the idea that industries exercise higher market power under protection.

Regarding the impact of tariffs, the results in column (2) suggest that reductions in ITR are associated with productivity gains while the effects of NTR are insignificantly different from zero. A one percentage point decline in ITR implies an increase in productivity growth rate of approximately 0.4%. This result favors the hypothesis that productivity gains are realized through the decline in input costs which increase domestic producers' access to foreign intermediate and capital goods. We note that our findings under the Roeger scenario are not biased by the assumption of CRS. As will be discussed further on, testing the coefficient of the scale parameter, β, resulting from the Harrison procedure, we do not reject CRS at a 5% significance level. However it is important to note that our findings might be understating the contribution of NTR to productivity growth. As noted in Section 3, by construction and due to the level of industry aggregation used to compute ITR, ITR for an industry j is likely to also include the NTR applied to that respective industry. Accordingly part of the effect of NTR will indirectly operate on efficiency growth through our ITR measure. In other words, the total effect of changes in NTR on productivity growth for an industry j is given by $(\widehat{\gamma_{NTR}} + (b_{jj} * \widehat{\gamma_{ITR}}))$. The magnitude of this total effect is thus dependent on the size of b_{jj}, the latter capturing the proportion by which an industry feeds itself with inputs. Examining the Quantec Input-Output table we find that the Motor Vehicle and the Communication Equipment industries are the largest self-feeding sectors, with inputs from the same industry accounting for 34% and 31% respectively of each industry's gross output, while the self-feeding rate for 21 of the remaining 26 industries is less than 15%.

Columns (3) & (5) report findings based on the Harrison technique applied to estimation equation (7), and columns (4) & (6) refer to the findings from estimating equation (4) in which we implement the Hall procedure. In this section we discuss the results when we do not control for changes in capacity utilization since accounting for such adjustments under

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both the Harrison and Hall procedures, does not significantly alter our findings.[11] We report the results from both these techniques given the previously noted advantages to each. On the one hand, the Harrison technique provides an estimation technique that does not require approximating an unobserved user cost of capital, r. Additionally it allows estimating the scale elasticity parameter, β, which is crucial to confirm the assumption of CRS that underlies the Roeger technique. On the other hand, the Hall technique has the advantage of estimating only one parameter, namely the markup, thus provides a higher level of estimation efficiency. Furthermore, the two estimation procedures model the control for capacity utilization differently providing additional robustness checks.

Columns (3) & (4) display the results from the Harrison and Hall estimation procedures, respectively, without controlling for the endogeneity of inputs. The estimates of both the markup and the impact of tariffs on efficiency growth are similar to those reported under the Roeger procedure. These results are further maintained in columns (5) & (6) in which we control for the endogeneity of inputs, dx_{jt}, using lagged values of capital, labor and intermediate inputs. The relevance of our instruments is confirmed by the Cragg-Donald test for weak instruments indicating that the bias in the IV estimator relative to the OLS bias is in the range of 10% to 20%. Furthermore the Hansen statistic confirms the validity of our IVs. With regard to the relationship between trade protection and the markup, different to columns (1), results when using the Harrison and Hall estimators imply insignificant coefficients on the interaction term reported in columns (3) to (6). Thus, controlling for endogeneity bias (stemming from the interaction between markups and tariffs), is key for why we find that the markup is significantly affected by changes in trade protection.

A notable feature of our findings is that the Roeger technique yields very similar results with respect to the effect of tariffs on efficiency growth to those reported by both the Harrison and Hall procedures. Moreover, the magnitude of the effect of the reductions in NTR on efficiency growth is stable across all three methodologies. Using the Roeger procedure, Edwards (2005b) also finds evidence in support of increased market power under higher levels of protection. Additionally, Fedderke et al (2006), using industry level data from 1970 to 1997, find that increased import penetration and export intensity serve to lower markups in South Africa. On the other hand, Aron & Muellbauer (2007) using aggregate annual time series data from 1971 to 2005 find evidence suggesting the slow adjustment of output prices in the short run. With the increase in trade openness, prices of imports and unit labor costs

[11] Tables 4 & 5 report the results in which we control for changes in capacity utilization.

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decrease resulting in a net effect of a rise in markups before the long run effect of increased competition feeds through. Finally, Aghion et al (2006) show that for the reform period, from 1995 to 2004, the Lerner index for most manufacturing industries did not decline. In light of this empirical evidence, there does not seem to be a consensus regarding the effect of changes in trade policy on the markup in SA.

To elaborate the extent of our findings and based on the results from our preferred specification in column (2), we compute the differential effect of the period reductions in NTR and ITR on TFP growth for both the Wearing Apparel and the Tobacco industries which witnessed the highest period cuts in NTR. With respect to the former industry, the reductions of 50.44 and 16.86 percentage points in NTR and ITR translate to an insignificant effect of NTR on productivity growth compared to an increase of 7.92% (-0.470*16.86) due to the cuts in ITR. Yet accounting for the previously mentioned bias in the way ITR are constructed, the 50.44 percentage point reduction in NTR increased the industry's productivity growth rate by 0.20% (-0.470*50.44*0.0084).[12] Regarding the Tobacco industry, the reductions in NTR and ITR of 45.44 and 6.18 percentage points resulted in an increase of 0.05% (-0.470*45.44*0.0023) and 2.9% (-0.470*6.18) in productivity growth rates, respectively.

A common concern in the empirical work that assess the effect of tariffs on productivity is the endogeneity of protection. In our results we address this concern by, firstly, employing lagged tariffs in our estimations. Secondly, we include industry fixed effects to control for the unobserved time invariant industry characteristics affecting, simultaneously, productivity and tariffs. Third, and as discussed in Section 3, the structure of the tariff schedule suggests that industry selectivity reductions or lobbying were somewhat limited during the reform period. Finally, we argue that SA's new government liberal trade policy position in 1994 was triggered by their plan of raising industrial efficiencies by curbing domestic monopoly power that had vested interests in the prevalent protectionist policies. Accordingly lower tariffs were applied to sectors with lower efficiency. Given our results where we find a negative coefficient on the tariff variable, solving for this endogeneity bias will serve to further increase the magnitude of the negative impact of tariffs on productivity growth, and in this respect will further enforce our findings that suggest that reduction in ITR induce increases in productivity growth.

In Table 3 we report the effect of changes in NTR on productivity growth and we do not control for the impact of changes in ITR on productivity or the effect of ETR on markups.

[12] Where b_{jj}=0.0084, i.e. 0.84% of the Wearing Apparel industry's gross output is supplied by its own industry. Similarly, b_{jj}=0.0023 for the Tobacco industry.

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This scenario produces results that are comparable to other findings in the literature. Our estimates suggest a significant negative relation between NTR and productivity growth. A one percentage point decline in tariffs translates to approximately 0.1% increase in productivity growth. This finding implies the plausible bias resulting from not controlling for ITR as productivity gains under this scenario are solely attributed to the competitive pressures from cheaper foreign imports of finished goods.

So far we discussed results from three different estimation techniques. We also show in Tables 4 & 5 that our findings are robust to controlling for capacity utilization. We perform additional robustness checks. Firstly, to further confirm the dominant effect of the reductions in ITR as a positive predictor of efficiency growth we repeat our exercise using ETR instead of NTR, the former being a measure for the effect of *net* competition. We find that the positive effect of ITR persists while the effect of competition, proxied by ETR, is insignificant. Secondly, we address the quality issue of the employed capital stock data, similar to Edwards & Golub (2004) we construct an alternative proxy for capital based on Harrigan (1999):

$$K_{jt} = \sum_{n=1}^{T} I_{j,t-n}(1-\sigma)^{n-1} \tag{2.17}$$

where I_{jt} is gross investment in sector j at time t deflated by the fixed investment deflator. We assume a useful life of capital good of ten years ($T = 10$) and a depreciation rate of 10% ($\sigma = 0.10$). Given that we have annual data on I_j from 1970, the computed capital stock under this procedure begins in 1979 (for our purposes we use the capital stock starting 1994). Our results are robust to this measure of capital stock.[13] Thirdly, we include trade flow variables in our estimations, namely: import penetration and export intensity ratios. Finally, to ensure that our results are not driven by the presence of outliers, we run our regressions excluding industries that witnessed the biggest reductions in tariff rates. Our results are confirmed under all robustness checks.

2.5 Conclusion

The South African manufacturing sector witnessed dramatic productivity growth between 1992 and 2000, coinciding with increases in trade openness. In this paper we examined the impact of tariff cuts on TFP growth exploiting industry level data from 1994 to 2004. We

[13] We do not present the results from these specification as they do not differ from those reported.

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identify two mechanisms through which tariff cuts can affect efficiency growth by distinguishing the differential effect of reductions in Nominal Tariff Rates, which capture the effect of increased foreign competition, and reductions in Intermediate Input Tariffs Rates, which capture the impact of imported technology. We carefully estimate the effect of tariffs on efficiency as we employ empirical estimation methodologies that separately control for the variation in markups attributed to changes in protection levels.

Our finding suggest that the efficiency difference between foreign and domestic inputs had a major effect on TFP gains where it is through the new and sophisticated technology embodied in the cheaply available imported input that trade openness positively affects efficiency growth. A one percentage point decline in ITR translates to 0.4% increase in efficiency growth compared to an insignificant effect of NTR. This finding is robust across all estimation procedures. It is also robust to controlling for: the endogeneity of inputs, tariffs and the interaction of tariffs and markups; changes in capacity utilization over the business cycle; to using an alternative capital stock series; and finally to employing Effective Tariff Rates (ETR) as an alternative proxy, to NTR, that captures the effect of increased import competition. Candidate explanations for the insignificant (or small) impact of increased competition on industry productivity growth are the previously outlined features of developing economies. The resource reallocation effect may not materialize as credit constraints may impede the expansion of efficient firms, while the lack of secondary markets for capital goods and the inflexibility of labor regulations can obstruct the exit of the inefficient ones. Additionally, the stringent labor market conditions regarding the hiring and firing of workers can also serve to block efficiency gains that can result from the impact of increased competition on reducing managerial inefficiencies. It is important to note that inflexible labor markets is a particular feature of the South African economy in which trade unions play a major role. Alternatively, and as outlined in Tybout (2000), the effect of openness on resource reallocation and on reducing managerial inefficiencies is more likely to have a static impact, thus may only affect productivity *levels* as opposed to affecting productivity *growth*. With regard to the effect of tariff reductions on markups, our findings suggest a decline in market power during the reform period, yet this result is not robust across all estimations.

The results give rise to some concerns with regard to the impact of trade reform in SA. There are foregone gains from the implemented free trade policies given that competitive pressures fail to translate to significant productivity gains or to confirmed reductions in markups. Investigating these issues in more detail would be an interesting avenue for future research.

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2.A Tables

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Table 1
Tariff Rates (%)

		Tariff on Final Output (NTR)					Tariff on Input (ITR)				
		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
ID	Industry	1993	1999	2004	Δ	$\widetilde{\Delta}$	1993	1999	2004	Δ	$\widetilde{\Delta}$
5	Food	22.94	13.53	11.19	(11.74)	(9.55)	9.75	5.28	4.38	(5.37)	(4.89)
6	Beverages	36.02	13.94	12.29	(23.73)	(17.45)	12.30	4.57	4.04	(8.26)	(7.35)
7	Tobacco	75.10	33.27	29.66	(45.44)	(25.95)	10.08	4.69	3.90	(6.18)	(5.61)
8	Textiles	47.76	29.57	16.53	(31.23)	(21.14)	19.08	10.04	6.22	(12.86)	(10.80)
9	Wearing Apparel	81.46	52.38	31.02	(50.44)	(27.80)	26.62	17.05	9.76	(16.86)	(13.32)
10	Leather & Leather Products	25.48	13.13	11.36	(14.12)	(11.25)	15.60	8.31	7.63	(7.97)	(6.89)
11	Footwear	46.77	25.34	22.40	(24.37)	(16.61)	16.37	8.19	7.69	(8.68)	(7.46)
12	Wood & Wood Products	17.82	8.93	8.67	(9.15)	(7.77)	6.52	3.09	3.43	(3.09)	(2.90)
13	Paper & Paper Products	12.38	7.11	6.46	(5.93)	(5.27)	6.80	3.29	3.30	(3.50)	(3.28)
14	Printing & Publishing	17.48	4.89	4.69	(12.79)	(10.89)	5.27	2.46	2.73	(2.54)	(2.41)
15	Coke & Refined Petroleum	12.98	4.60	3.37	(9.61)	(8.51)	1.75	0.57	0.82	(0.93)	(0.91)
16	Basic Chemicals	8.42	1.97	1.67	(6.75)	(6.23)	4.27	1.41	1.25	(3.02)	(2.90)
17	Other Chemicals	16.93	5.09	4.35	(12.59)	(10.76)	7.13	2.93	2.39	(4.74)	(4.43)
18	Rubber Products	21.10	12.53	10.57	(10.53)	(8.70)	6.42	3.17	2.23	(4.18)	(3.93)
19	Plastic Products	22.61	12.00	9.65	(12.96)	(10.57)	5.74	2.22	2.27	(3.47)	(3.28)
20	Glass & Glass Products	19.05	7.51	7.31	(11.74)	(9.86)	5.14	2.10	2.16	(2.98)	(2.83)
21	Non-Metallic Minerals	16.61	5.31	5.57	(11.04)	(9.47)	3.53	1.26	1.42	(2.11)	(2.04)
22	Basic Iron & Steel	9.69	4.32	3.89	(5.80)	(5.29)	4.58	1.76	1.58	(2.99)	(2.86)
23	Basic Non Ferrous Metal	9.38	2.58	1.98	(7.40)	(6.77)	4.49	1.15	0.93	(3.56)	(3.41)
24	Metal Products	20.57	8.05	7.84	(12.73)	(10.56)	5.84	2.17	2.59	(3.25)	(3.07)
25	Machinery & Equipment	12.98	3.99	3.44	(9.55)	(8.45)	6.79	2.74	2.54	(4.25)	(3.98)
26	Electrical Machinery	21.16	8.12	7.15	(14.01)	(11.56)	7.89	3.15	2.91	(4.98)	(4.62)
27	Communication Equipment	24.77	3.51	2.73	(22.04)	(17.66)	11.92	2.64	2.42	(9.50)	(8.48)
28	Prof. & Sci. Equipment	13.79	0.33	0.33	(13.46)	(11.83)	8.96	3.61	2.34	(6.62)	(6.08)
29	Motor Vehicles	27.83	18.28	14.64	(13.19)	(10.32)	16.00	10.47	9.66	(6.34)	(5.47)
30	Other Transport Equip.	13.16	1.47	0.85	(12.32)	(10.88)	5.28	1.48	1.72	(3.56)	(3.38)
31	Furniture	32.48	17.60	17.37	(15.12)	(11.41)	9.15	6.18	5.52	(3.63)	(3.32)
32	Other Manufacturing	27.39	6.61	5.82	(21.56)	(16.93)	4.43	2.71	2.45	(1.98)	(1.90)
	Mean	25.50	11.64	9.39	(16.12)	(12.12)	8.85	4.24	3.58	(5.26)	(4.71)
	Standard Deviation	17.93	11.53	7.98	10.79	5.67	5.56	3.64	2.49	3.50	2.79

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

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2.B Figures

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Table 2
The Effect of Reductions in NTR and ITR on Productivity Growth
Dependent Variable: Growth in Real Output

	(1)	(2)	(3)	(4)	(5)	(6)
Scale Parameter (β)			1.057**		1.164**	
			(0.035)		(0.087)	
Markup (μ)	1.144**		1.184**	1.179**	1.206**	1.257**
	(0.027)		(0.035)	(0.035)	(0.086)	(0.084)
Lag ETR*dx_{jt} (μ_{ETR})	0.103*		0.015	0.004	0.039	-0.057
	(0.035)		(0.032)	(0.039)	(0.057)	(0.074)
Lag NTR		0.025	0.018	0.027	0.028	0.026
		(0.063)	(0.063)	(0.062)	(0.061)	(0.060)
Lag ITR		-0.470*	-0.432*	-0.463*	-0.450*	-0.439*
		(0.193)	(0.189)	(0.190)	(0.191)	(0.182)
Observations	308	308	308	308	308	308
Adjusted R-squared	0.96	0.12	0.93	0.93	0.93	0.93
Estimation Procedure	Roeger	Roeger	Harrison	Hall	Harrison	Hall
IV	No	No	No	No	Yes	Yes
Test of IV Identification:						
Hansen Test (p-value)					0.147	0.151
Cragg-Donald (CD) F-Statistic					6.345	8.928
CD Critical Value (% relative bias)					8.78(10%)	9.08(10%)
CD Critical Value (% relative bias)					5.91(20%)	6.46(20%)
Test of Coefficients (p-value):						
Constant Return Scale, $H_0 : \beta = 1$			0.110		0.061	
Market Power, $H_0 : \mu = 1$	0.000		0.000	0.000	0.016	0.002

Table 3
The Effect of Reductions in NTR on Productivity Growth
Dependent Variable: Growth in Real Output

	(1)	(2)	(3)	(4)	(5)	(6)
Scale Parameter			1.055**		1.166**	
			(0.036)		(0.089)	
Markup	1.170**		1.189**	1.182**	1.236**	1.254**
	(0.025)		(0.033)	(0.033)	(0.081)	(0.074)
Lag NTR		-0.098*	-0.100*	-0.098*	-0.096*	-0.096*
		(0.040)	(0.038)	(0.040)	(0.039)	(0.038)
Observations	308	308	308	308	308	308
Adjusted R-squared	0.96	0.11	0.93	0.93	0.93	0.93
Estimation Procedure	Roeger	Roeger	Harrison	Hall	Harrison	Hall
IV	No	No	No	No	Yes	Yes
Test of IV Identification:						
Hansen Test (p-value)					0.143	0.143
Cragg-Donald (CD) F-Statistic					5.915	9.738
CD Critical Value (% relative bias)					8.78(10%)	13.91(5%)
CD Critical Value (% relative bias)					5.91(20%)	9.08(10%)
Test of Coefficients (p-value):						
Constant Return Scale, $H_0 : \beta = 1$			0.128		0.062	
Market Power, $H_0 : \mu = 1$	0.000		0.000	0.000	0.003	0.000

Robust standard errors in parentheses, + significant at 10%;*significant at 5%; ** significant at 1%.
All regressions include time and industry dummies.

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Table 4
The Effect of Reductions in NTR and ITR on Productivity Growth
(with capacity utilization control)
Dependent Variable: Growth in Real Output

	(1)	(2)	(3)	(4)
Scale Parameter	1.028**		1.072**	
	(0.028)		(0.059)	
Markup	1.111**	1.107**	1.126**	1.167**
	(0.045)	(0.032)	(0.090)	(0.077)
Lag ETR*dx_{jt}	0.023	0.005	0.042	-0.041
	(0.023)	(0.037)	(0.031)	(0.065)
Lag NTR	0.019	0.013	0.016	0.011
	(0.062)	(0.064)	(0.059)	(0.060)
Lag ITR	-0.419*	-0.403*	-0.389*	-0.382*
	(0.185)	(0.191)	(0.181)	(0.180)
Observations	308	308	308	308
Adjusted R-squared	0.93	0.93	0.93	0.93
Estimation Procedure	Harrison	Hall	Harrison	Hall
IV	No	No	Yes	Yes
Tests of IV Identification:				
Hansen Test (p-value)			0.576	0.145
Cragg-Donald (CD) F-Statistic			4.959	9.448
CD Critical value (% relative bias)			5.91(20%)	9.08(10%)
CD Critical value (% relative bias)			4.79(30%)	6.46(20%)
Test of Coefficients (p-value):				
Constant Return Scale, $H_0 : \beta = 1$	0.307		0.221	
Market Power, $H_0 : \mu = 1$	0.013	0.009	0.160	0.030

Table 5
The Effect of Reductions in NTR on Productivity Growth
(with capacity utilization control)
Dependent Variable: Growth in Real Output

	(1)	(2)	(3)	(4)
Scale Parameter	1.030**		1.075**	
	(0.028)		(0.059)	
Markup	1.120**	1.110**	1.162**	1.164**
	(0.043)	(0.029)	(0.085)	(0.069)
Lag NTR	-0.096*	-0.096*	-0.094*	-0.095*
	(0.039)	(0.036)	(0.037)	(0.034)
Observations	308	308	308	308
Adjusted R-squared	0.93	0.93	0.93	0.93
Estimation Procedure	Harrison	Hall	Harrison	Hall
IV	No	No	Yes	Yes
Tests of IV Identification:				
Hansen Test (p-value)			0.374	0.136
Cragg-Donald (CD) F-Statistic			5.025	10.27
CD Critical value (% relative bias)			5.91(20%)	13.91(5%)
CD Critical value (% relative bias)			4.79(30%)	9.08(10%)
Test of Coefficients (p-value):				
Constant Return Scale, $H_0 : \beta = 1$	0.296		0.207	
Market Power, $H_0 : \mu = 1$	0.000	0.000	0.057	0.018

Robust standard errors in parentheses, + significant at 10%;*significant at 5%; ** significant at 1%.
All regressions include time and industry dummies.

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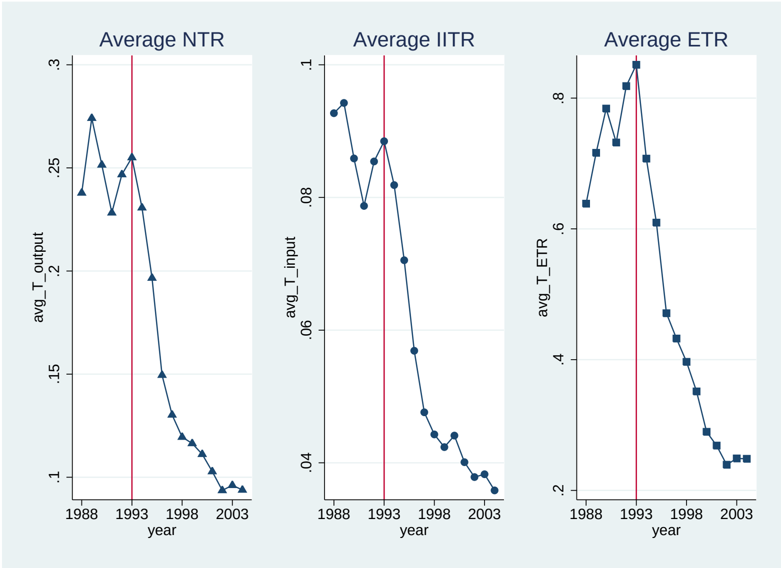

Figure 1: Yearly Simple Average of Tariff Rates

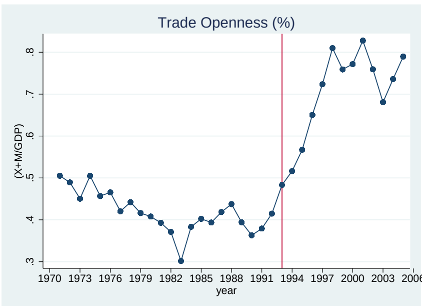

Figure 2: Trade Openness (%) (Export + Import)/GDP

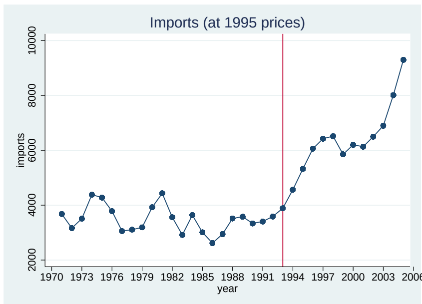

Figure 3: Real Imports (1995 local prices)

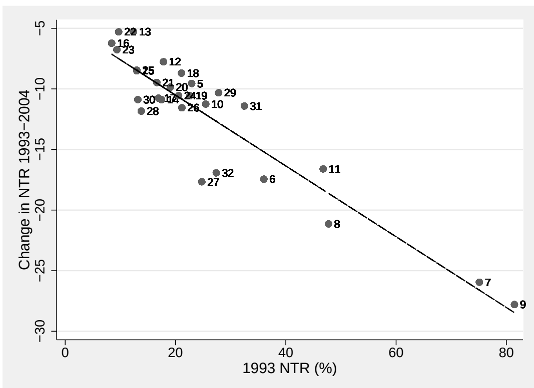

Figure 4: Tariffs on Final Output (NTR)

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Chapter 3

Heterogenous Trade Barrier Effects: What Can We Learn from a Disaggregated Gravity Model?

3.1 Introduction

It is common practice in the gravity literature is estimate the elasticity of trade flows with respect to trade barriers using aggregate data. If the sectoral responses to changes in trade barriers are heterogenous, results from aggregate estimates may be biased. In this paper, using bilateral trade flow data between South Africa (SA) and approximately 160 trading partners for 28 manufacturing industries and covering the period from 1995 to 2004, we test two hypotheses that predict heterogenous impact of trade barriers. We examine two sources of heterogeneity: variations in the degree of firm heterogeneity in size across sectors, and differences in the export size of the trading partners. We investigate whether the negative impact of trade barriers on trade flows is higher for sectors with more homogenous firms and is lower for larger trading partners. The results confirm these predictions. We find that the negative elasticity of SA's *exports* to trade barriers is higher for sectors with a larger level of firm homogeneity, while the elasticity of *imports* to trade barriers is lower for larger trading partners.

Anderson and Wincoop (2004) highlight the bias that stems from estimating trade costs using aggregate data when trade costs vary at the disaggregated level. They note that "while aggregation bias can theoretically arise in many different ways, little is known about the empirical magnitude of the bias. One obvious recommendation is to disaggregate." We highlight some of the documented sources of aggregation bias. The first emerges from the assumption imposed on the elasticity of substitution across varieties. The underlying premise in estima-

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tions based on aggregate data is that there is a common elasticity of substitution for all goods, an assumption unlikely to hold in the real world. Secondly, aggregation bias can be due to zero bilateral trade flows between country pairs. Sectoral data reveal a large number of zero observations in cross border pairs in comparison to aggregate data. Ignoring these zeros leads to an over estimation of the impact of trade barriers. Finally, industry structural differences can be an additional source of bias. As noted in Haveman et al (2003) "using disaggregated data mitigates specification error due to structural differences in determinants of trade flows across commodities."

In the first part of this study we test the hypothesis that the impact of trade barriers on SA's *exports* is larger in more homogenous sectors. Chaney (2008), who extends the Krugman (1980) model to allow for firm heterogeneity in productivity and fixed costs of exporting, shows that the negative impact of trade barriers on trade flows is magnified for more *homogenous sectors*, the latter defined as sectors where large productive firms represent a smaller fraction of firms, as opposed to *heterogenous sectors* in which large productive firms account for a larger fraction of firms. In the model, reductions in trade barriers raise each firm's exports (intensive margin), moreover new firms start exporting (extensive margin). This adjustment in the extensive margin is quantitatively important. In more homogenous sectors, large productive firms represent a smaller fraction of firms. Given a change in variable trade costs, the productivity threshold moves in a region where most of the mass of firms exist. In those sectors, aggregate exports are sensitive to changes in trade costs because many firms exit and enter as variable costs fluctuate. To test this prediction we employ a gravity model using SA's bilateral exports as our dependent variable. We use the C5% Index derived in Fedderke & Naumann (2008) as our proxy for sectoral homogeneity. This measure of industry concentration denotes the cumulative percentage of output accounted for by the dominant 5% of firms. A higher value indicates a more concentrated and, hence, a less competitive sector. We believe that this measure represents a good proxy for sectoral homogeneity given that the distribution of firm size is shaped by the degree of competition.

In the second part of the paper we test whether the elasticity of SA's *imports* to trade barriers depends on the size of the exporting partner. In models with fixed costs of exporting (Romer, 1994), the impact of trade barriers is inversely related to the size of the market in the trading partner country. Haveman et al (2003) show that in the presence of country specific fixed costs, trade is compressed into fewer partners than would otherwise occur in the absence of trade barriers. Testing this *compression effect* for multilateral tariffs, they find that higher multilateral tariffs shift trade towards importing from larger potential exporters. They argue that this can be due to the objective of minimizing fixed costs of trading that is related to the number of countries with which the importer trades, a goal that may dominate any preference for variety. In line with their work, we use the trading partners' world exports to capture its export potential. Using SA's bilateral imports as our dependent variable in the gravity model, we test for the asymmetric effect of trade barriers attributed to the heterogeneity in partners' export size.

To examine the above mentioned hypotheses we exploit annual trade flow data for 28 manufacturing industries in South Africa covering the post apartheid period from 1995 to 2004. We find that the elasticity of SA's *exports* to trade barriers is higher for more homogenous sectors, while the elasticity of SA's *imports* to trade barriers is lower for trading partners with larger export potential. We note that our findings are robust across two estimation procedures, the Tobit and the Heckman Sample Selection Model.

This paper is divided into four sections. In the next section we present the empirical methodology. Section 3 describes the data used in this paper. Section 4 discusses our results, Section 5 concludes.

3.2 Methodology

The simplest form of the gravity model states that bilateral trade flows between two countries is positively related to the product of their GDPs and inversely proportional to their distance. The model extends to include other observable characteristics that proxy trade barriers between the countries (e.g. common language, geographic isolation, common border). Lacking a theoretical foundation, this model was firstly applied to international trade by Tinbergen(1962), Poyhonen(1963) and Linneman(1966). The model performs well empirically, providing sensible parameter estimates and explaining large parts of bilateral trade. Economic theory caught up with the empirical evidence giving rise to number of theoretical models that predict the gravity model (see Feenstra 2004, Anderson and Wincoop 2003 for recent surveys). The baseline specification of the gravity model we adopt in this study is as follows:

$$Trade_{jtk} = \alpha_1 GDP_{jt} + \alpha_2 GDPC_{jt} + \beta_1 H_j + \alpha_t + \alpha_k + \eta_{jtk} \tag{3.1}$$

where $Trade_{jtk}$ is the logarithm of trade of good k in year t between SA and trading partner j. Depending on the question studied, we employ exports or imports as the dependent variable.

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GDP_{jt} is the logarithm of GDP of trading partner j in year t. This variable captures the market size effect of the partner country. $GDPC_{jt}$ is the logarithm of GDP Per Capita in country j to account for differences in standards of living. H_j is a matrix of trade barriers relating SA to country j. We focus on three barriers that are shown robust in the literature namely: distance between countries, a dummy if they share a common language and a dummy if j is landlocked. We also include industry dummies, α_k, to control for the time invariant industry specific characteristics such as the elasticity of substitution. Moreover we include time dummies, α_t, to capture period related macroeconomic factors that are common across sectors. Given our focus on SA's trade patterns, we do not control for its own GDP or GDPC in our regressions. The impact of these variables are captured by the time controls.

A common challenge in the empirical estimation of the gravity model is how to deal with zero bilateral trade flows. A wide-spread practice is to perform OLS estimation on the positive trade flows, ignoring the zero observations. If zero flows do not occur randomly, the latter technique will deliver biased estimates and the effect of trade barriers, being a plausible force underlying this missing trade, is underestimated. Acknowledging that the zero dependent variables carry information that ought to be included in the estimation calls for other econometric procedures that specifically address censored data. Given that trade values are bounded from below by zero, the appropriate estimation procedure is a Tobit model. This specification is used in a number of works (e.g. Rose(2004), Soloaga and Winters (2001), Anderson and Marcouiller(2002) and Venables(2004)). An alternative technique often employed in the literature is the two-stage Heckman(1979) Sample Selection Model. The first stage entails a probit equation that determines the decision of a country pair to trade or not. The second stage involves modeling the observed positive trade flows, augmenting the regression with the inverse Mills ratio from the first stage to account for selection bias. Due to the lack of a theoretically motivated exclusion restriction in the former stage, both stages in the Heckman procedure have the same specification. Accordingly the non linearity of the Mills ratio is the only source of identification. In this paper we report findings from both the latter techniques. We consider the Tobit results to be our preferred specification in light of the aforementioned weakness of identification in the Heckman procedure. Moreover, we believe that a Heckman specification may be more relevant for finer data (e.g. firm level data) where the decision to trade is more sophisticated and is more likely to be independent from the decision regarding the volume of trade.

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3.2.1 Hypothesis 1: Heterogenous Trade Barrier Effects due to Differences in the Level of Sectoral Firm Homogeneity

In this section we outline the empirical specification to test whether the negative impact of trade barriers on exports is exacerbated for more homogenous sectors. As previously noted, our measure of the level of firm homogeneity within a sector is captured by the C5% Index developed in Fedderke et al (2008). A higher value indicates a more homogenous sector. Accordingly our empirical model is as follows:

$$Exports_{jtk} = \alpha_1 GDP_{jt} + \alpha_2 GDPC_{jt} + \beta_1 H_j + \beta_2 (H_j * CIndex_{tk}) + \alpha_k + \alpha_t + \eta \qquad (3.2)$$

where $Exports_{jtk}$ is ln exports of SA of industry k to country j at year t, estimated in USD and deflated by the CPI for the USA. $CIndex_{tk}$ is $ln(1 + C5\%Index)$. Based on Chaney (2008) we expect the coefficient on the trade barrier and the interaction term to have the *same* sign.

3.2.2 Hypothesis 2: Heterogenous Trade Barrier Effects due to Differences in the Level of Trading Partners' Export Potential

To test our second hypothesis, if the negative impact of trade barriers on imports is reduced for trading partners with larger export potential. Our empirical specification is:

$$Imports_{jtk} = \alpha_1 GDP_{jt} + \alpha_2 GDPC_{jt} + \beta_1 H_j + \beta_2 (H_j * Size_{tk}) + \beta_3 \tau_{tk} + \beta_4 (\tau_{tk} * Size_{tk}) + \alpha_k + \alpha_t + \eta \qquad (3.3)$$

where $Imports_{jtk}$ is the logarithm of imports of SA of industry k from country j at year t, estimated in USD and deflated by the CPI for the USA. $Size_{jt}$ is equal to the logarithm of real world exports of country j in year t. τ_{tk} is the log of $(1 + NTR_{tk})$ where NTR_{tk} is Nominal Tariff Rates applied by SA on imports of good k in year t. Our hypothesis suggests that the coefficient on the trade barriers and the interaction term have *opposite* sign.

We note the concern of simultaneity bias in the above estimation with respect to the tariff variable. Higher levels of imports may be driving higher tariff rates. We dismiss this concern in our estimation in light of SA's new government liberal trade policy position in 1994 which was triggered by their plan of raising industrial efficiencies by curbing domestic monopoly power that had vested interests in the prevalent protectionist policies (Bell 1997). Accordingly the trade policy, as stated, is independent of the level of industry trading activities. Moreover, we

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believe that industry fixed effects in our estimation control for the unobserved time invariant industry characteristics affecting, simultaneously, imports and tariffs.

3.3 The Data

We study bilateral trade flow data for the manufacturing sector provided by Quantec. The data are disaggregated to 28 industries at the SIC-3 digit level and covers trade with 160 partners (excluding countries with population size less than one million inhabitants). Our period starts in 1995 after the first post-apartheid government was democratically elected. Our starting year ensures that all sanctions had been lifted, furthermore it restricts our analysis to the particular political time frame of the first post apartheid regime.

As previously mentioned, to test the first hypothesis we use the C5% Index from Fedderke et al. (2008) as our measure of the level of firm heterogeneity within a sector. The index is computed as the percentage of output accounted for by the dominant 5% of firms. A higher value indicates a more concentrated sector with a lower level of competition. Given that the size distribution of firms is shaped by the degree of competition we believe that this proxy is a fair representation of sectoral homogeneity as defined in Chaney (2008). The data are provided for 23 of the 28 sectors. The C5% Index shows considerable variation across the 23 sectors ranging from a maximum of 85% for the motor vehicle industry in 1996 to a minimum of 48% for the textile industry. The average index decreased from 68% in 1996 to 54% in 2001 while the coefficient of variation increased from 15% to 31%, respectively. Given this data limitation to only two years, we interpolate the concentration index for 1998 as the average of the 1996 and 2001 estimates. We use 1996 estimates for the years the from 1995 to 1997, the 1998 figures also for the year 1999, and we use 2001 figures for the years from 2000 to 2004.[1]

Regarding the second hypothesis we use tariff data to capture the impact of trade policy on SA's imports. We employ the Nominal Tariff Rates estimated by Edwards (2005). The data are provided for the entire period and covers our 28 manufacturing industries. Confining our analysis to using data from the mid nineties provides a consistent tariff series as the *ad valorem* equivalent of the Non Tariff Barriers are estimated for this period. Noteworthy that for the period under study and given our focus on the manufacturing sector, all trading partners are subject to Most Favored Nation (MFN) tariff rates, accordingly tariffs vary only

[1] Our results are also robust to confining our regressions to the years 1996, 1998 and 2001.

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by sector but are common across all countries.[2]

With respect to the remaining variables, we use figures for GDP and GDP per Capita from the Penn World Table. Data on country pair trade barriers namely: distance, common official language and geographic isolation (landlocked), is obtained from the Centre D'Etudes Prospectives Et D'Informations Internationales (CEPII).

3.4 Results

Table 1 shows the estimation of the model in equation (2). Columns (1) to (4) report results from our preferred specification, the Tobit estimator, while columns (5) to (9) display findings from the Heckman Sample Selection Model, where column (9) shows the first stage results for the selection decison.

Results for the baseline scenario in column (1) are in line with general gravity model predictions. Both the size and the living standard of the trading partner have a positive impact on SA's exports, as implied by the positive coefficients on the GDP and the GDP per Capita variables. The elasticity estimate of 1.9 suggests that a one percent increase in partner GDP raises SA's exports by 1.9%. Moreover, both distance and geographic isolation, the latter depicted by a trading partner dummy if landlocked, have a negative impact on trade, while sharing a common official language with the trading partner increases SA's exports. The results concerning the interaction parameter reported in columns (2) to (4) support the theoretical predictions from Chaney (2008). A higher concentration index (thus a more heterogenous sector) magnifies the impact of trade barriers on trade flows. This is confirmed by the sign on the coefficient of the interaction between the trade barriers and our proxy for industry homogeneity, which takes the same sign as that on the relevant barrier. To quantify the magnitude of our findings, we compare two sectors, the motor vehicle industry with a period average concentration level of 81%, to the textile industry with a lower value of 41%. Results for the distance variable suggests that a one percent increase in distance between SA and its trading partner reduces SA's exports by 2.9% [-2.774+ (-0.267*ln(1+0.81))] for the former industry and 2.8% [-2.774+ (-0.267*ln(1+0.41))] for the latter. These estimates

[2] We note that SA's Free Trade Agreements are irrelevant to our sample. September 2000 was the implementation of the South African Development Community (SADC) Free Trade Protocol with little impact on the manufacturing sector. January 2000 was the implementation of the SA-EU Trade, Development & Cooperation Agreement (TDCA) yet ratification and enforcement of agreement was only in 2004. Finally, December 2004 marked the Preferential Trade Agreement between the South African Customs Union (SACU) & MERCOSUR which does not impact our sample period.

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suggest that the asymmetric impact of distance on trade flows due to variations in sectoral homogeneity, although significant, is negligible in magnitude.

In columns (5) to (9) we display the matching results from the Heckman procedure. Column (9) shows the estimates for the first-stage selection equation which models the decision of an industry to export or not. Regarding the second-stage results, displayed in columns (5) to (8), our estimates for the the distance and landlocked variables are significant and confirm our earlier findings of an amplified impact of trade barriers for more homogenous sectors, while our results regarding the language barrier are not robust.

Table 2 displays the results from estimation equation (3), for testing the second hypothesis regarding the compression effect of trade barriers. Again, columns (1) to (5) display results from the Tobit estimation and columns (6) to (11) from Heckman. Given that the dependent variable in these regressions is imports, we include nominal tariff rates as an additional trade barrier. One common concern is the simultaneity bias with respect to tariffs. Higher levels of imports may lead to higher tariff rates. We argue that this is not the case for SA given the new government's liberal trade policy position which was triggered by their plan of raising industrial efficiencies by curbing domestic monopoly power that had vested interests in the prevalent protectionist policies (Bell 1997). Hence the trade policy, as stated, can meaningfully be assumed to be independent of the level of industry trading activity. Moreover, the industry fixed effects in our estimations serve to reduce endogeneity concerns as they control for the unobserved time invariant industry characteristics affecting, simultaneously, imports and tariffs.

As in Table (1), results in Table (2) column (1) are in line with the gravity model predictions where factors such as the trading partner's GDP and GDPC, and sharing a common official language with SA increase imports, while distance and being landlocked reduce trade. With regard to the tariffs barrier, the negative coefficient indicates that higher protection decreases imports, yet this effect is not significant. This may be due to the fact that MFN tariffs, being common across all partners, serve better to explain aggregate industry imports as opposed to bilateral imports. We note to the results from the Heckman selection equation in column (11) which suggest that tariffs are a significant predicator of an industry's decision to import. With respect to our findings regarding the asymmetric impact of trade barriers due to variations in partner export potential, results in columns (2) to (5) highlight the predicted compression effect as the impact of trade barriers on imports is smaller for partners with larger world export markets. This is confirmed by the coefficient on the interaction

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term which takes a sign opposite to that on the trade barrier. Our findings are robust for all trade barriers and under both techniques. Based on our findings in column (2) we quantify the extent of our estimates and compare the responsiveness of SA's imports to tariff changes across the USA and Mauritius, two trading partners with very different export market share. Our results suggest that a one percent rise in tariffs rates *increases* exports from the USA to SA by 2.2% [-13.247+ (0.566*27.27)] while it *reduces* exports from Mauritius to SA by 1.2% [-13.247+ (0.566*21.27)]. These findings are in line with Haveman et al (2003) who argue that the compression effect of multilateral tariffs can be due to the objective of minimizing fixed costs of trading that is related to the number of countries with which the importer trades, a desire that may dominate any preference for variety. Almost all results are robust under the Heckman procedure, reported in columns (7) to (10). We note that despite the consistency in the sign of the coefficient across both techniques, the magnitude of the estimates considerably differ.

3.5 Conclusion

The gravity model is commonly estimated such that the elasticity of trade flows with respect to trade barriers is equal across the different sectors and regions. This is the underlying assumption when gravity estimations are implemented using aggregate trade flow data. Supported by economic theory, in this paper we examine two sources of heterogeneity in trade barriers effects attributed to; (1) variations in sectoral level of firm heterogeneity, and (2) differences in trading partner export size.

Using annual bilateral data on exports and imports for South Africa with 160 trading partner, for 28 manufacturing industries and covering the period from 1995 to 2004, we test two hypotheses that predict a heterogenous effect of trade barriers on trade flows. Firstly, we examine if variations in the level of firm heternogeniety across sectors, measured by sectoral concentration ratio, implies asymmetric trade barrier effects. Secondly, we test if the heterogeneity in the impact of trade barriers can be attributed to differences in the size of the trading partners, the latter measured by the trading partner's export potential.

The signs on our estimated coefficients suggest that the negative elasticity of SA's *exports* to trade barriers is higher for more homogenous sectors. This result provides support for international trade models with heterogeneous firms as opposed to those assuming representative firms. Moreover, we find that the elasticity of SA's *imports* to trade barriers is lower

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for trading partners with larger export potential. This evidence highlights the presence of country specific fixed costs related to the number of countries with which the importer (SA) trades, and the desire to minimize these costs which may dominate any preference for variety. This can help explain the behavior of some developing countries who focus on exporting a variety of products from a few industries as opposed to diversifying their product space, a strategy that makes them big exporters in a small number of industries rather than small exporters in a big number of industries (Haveman et al. 2003).

Finally, we note that our use of disaggregated industry level data are important not only because it provides information on sources of asymmetric trade barriers effects, but also because it allows controlling for the aggregation bias that stems from the misrepresentation of zero bilateral trade flows in aggregate data. Additionally, it allows for industry fixed effects, hence controls for industry specific elasticity of substitution, a reality commonly overlooked in gravity estimations. We strongly encourage more empirical studies, particularly employing multinational data, to further support our findings.

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3.A Tables

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Table 1: Heterogeneity due to Industry Homogeneity
Dependent Variable: Ln(Exports)

	Tobit Regressions				Heckman Regressions				
					Truncated Regressions				Probit
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
GDP	1.932**	1.932**	1.933**	1.933**	1.210**	1.210**	1.209**	1.215**	0.323**
	(0.018)	(0.018)	(0.018)	(0.018)	(0.023)	(0.023)	(0.023)	(0.023)	(0.005)
GDPC	0.504**	0.504**	0.504**	0.504**	0.567**	0.567**	0.567**	0.568**	0.042**
	(0.039)	(0.039)	(0.039)	(0.039)	(0.019)	(0.019)	(0.019)	(0.019)	(0.008)
Distance	-2.901**	-2.774**	-2.901**	-2.902**	-2.780**	-2.748**	-2.778**	-2.789**	-0.302**
	(0.086)	(0.097)	(0.086)	(0.086)	(0.042)	(0.046)	(0.042)	(0.042)	(0.011)
Dist.*Con		-0.267**				-0.068+			
		(0.096)				(0.041)			
Language	2.050**	2.050**	1.353**	2.050**	1.526**	1.526**	2.000**	1.531**	0.338**
	(0.074)	(0.074)	(0.367)	(0.074)	(0.043)	(0.043)	(0.167)	(0.043)	(0.018)
Lang.*Con			1.467+				-1.003**		
			(0.765)				(0.341)		
Landlocked	-3.493**	-3.493**	-3.493**	-1.743**	-1.577**	-1.577**	-1.573**	-0.565**	-0.596**
	(0.108)	(0.108)	(0.108)	(0.541)	(0.066)	(0.066)	(0.066)	(0.192)	(0.019)
Llocked*Con				-3.691**				-2.176**	
				(1.131)				(0.387)	
Observations	37796	37796	37796	37796	37796	37796	37796	37796	37796

Robust standard errors in parentheses, + significant at 10%;*significant at 5%; ** significant at 1%.
All regression include Time and Industry Dummies.

Table 2: Heterogeneity due to Trading Partner Export Size
Dependent Variable: Ln(Imports)

	Tobit Regressions					Heckman Regressions					
						Truncated Regressions					Probit
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
GDP	2.594**	1.347**	2.677**	2.580**	2.533**	1.807**	1.202**	1.857**	1.856**	1.806**	0.440**
	(0.018)	(0.046)	(0.021)	(0.018)	(0.023)	(0.025)	(0.037)	(0.026)	(0.028)	(0.026)	(0.005)
GDPC	1.578**	0.704**	1.596**	1.344**	1.532**	1.353**	0.883**	1.369**	1.260**	1.351**	0.168**
	(0.035)	(0.049)	(0.035)	(0.037)	(0.037)	(0.024)	(0.031)	(0.024)	(0.027)	(0.024)	(0.008)
Distance	-1.550**	-4.420**	-1.550**	-1.628**	-1.545**	-1.130**	-2.764**	-1.127**	-1.190**	-1.130**	-0.219**
	(0.066)	(0.126)	(0.065)	(0.065)	(0.066)	(0.040)	(0.077)	(0.040)	(0.044)	(0.040)	(0.011)
Dist*Export		0.145**					0.079**				
		(0.005)					(0.003)				
Language	1.166**	1.262**	5.626**	1.319**	1.172**	0.831**	0.918**	3.417**	0.928**	0.831**	0.272**
	(0.071)	(0.071)	(0.647)	(0.071)	(0.071)	(0.046)	(0.050)	(0.350)	(0.050)	(0.046)	(0.017)
Lang.*Export			-0.200**					-0.113**			
			(0.028)					(0.015)			
Landlocked	-0.819**	-0.805**	-0.826**	-17.617**	-0.817**	-0.071	-0.143*	-0.076	-8.709**	-0.071	-0.248**
	(0.102)	(0.100)	(0.102)	(0.917)	(0.102)	(0.060)	(0.065)	(0.061)	(0.537)	(0.060)	(0.020)
Llocked*Export				0.776**					0.389**		
				(0.041)					(0.024)		
Tariff	-0.773	-0.73	-0.783	-0.756	-13.247**	-0.011	-0.046	-0.037	-0.055	-0.460	-0.710*
	(1.425)	(1.407)	(1.423)	(1.419)	(3.454)	(0.922)	(0.997)	(0.928)	(1.007)	(1.939)	(0.350)
Tariff*Export					0.566**					0.020	
					(0.135)					(0.076)	
Observations	42392	42392	42392	42392	42392	42392	42392	42392	42392	42392	42392

Robust standard errors in parentheses, + significant at 10%;*significant at 5%; ** significant at 1%.
All regression include Time and Industry Dummies.