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Gradual Trade Liberalization and Firm Performance in Ethiopia*

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Abstract

We use firm-level data for the Ethiopian manufacturing sector matched with commodity-level data on tariffs to examine the effect of trade liberalization on firm performance during the 1997-2005 period. We find relatively large positive effects of tariff reductions on total factor productivity, a result that is robust to treating tariffs as endogenous, and to various generalizations of the baseline model. This effect is primarily driven by mechanisms operating at high tariff levels, suggesting that excessive tariff levels may be particularly distortionary. We find some evidence that the reduction of tariffs has resulted in smaller and more capital-intensive domestic firms. We note that these effects are consistent with the hypothesis that the trade liberalization has increased competition in the domestic market. We find no significant effect of the trade liberalization on entry or exit rates.

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1 Introduction

Since the 1980s most countries in Sub-Saharan Africa have moved away from inward-looking development strategies, as a reaction to the failure of previous import substitution policies. The reforms were generally undertaken within the framework of structural adjustment programs under the auspices of the international financial institutions. Reforming trade policy, an important component of these programs, included import liberalization through tariff reduction and the removal of non-tariff barriers. Unlike other regions, there exists little empirical evidence on how trade reforms have impacted firm performance in Sub-Saharan Africa. In this paper we match firm-level panel data with commodity-level disaggregated data on imports and tariffs to investigate how trade reforms have affected manufacturing firms in one of the world's least developed countries: Ethiopia.

Several hypotheses exist about how increased openness to trade affects firm productivity. According to the import discipline hypothesis trade liberalization increases competition from imported goods, which forces domestic producers to improve their efficiency, by reducing managerial slack and using inputs more efficiently for example (Nishimizu and Robinson, 1984; Holmes and Schmitz, 2001). Increased competitive pressure may also lead to further exploitation of economies of scale, to the extent that the reduction in the market power of domestic firms forces them to expand output and move down the cost curve (Krugman, 1979; Helpman and Krugman, 1985). Import competition may have a re-allocation effect. Reduced protection lowers domestic prices, forcing high cost producers to exit the market while the efficient ones survive. This industry-rationalization may increase industry productivity by reallocation of resources from less efficient to more efficient producers, even with unchanged within-plant productivity (see, Roberts and Tybout, 1991; and Rodrik, 1992). Increasing integration into the world economy may benefit productivity because of other mechanisms than increased competition. Access to cheaper and better intermediate inputs, access to global finance, and exposure to new goods and new methods of production can be sources of increased productivity (Dornbush, 1992). Endogenous growth-trade theorists have formulated models in which trade opening contributes to economic growth, by increasing diffusion of knowledge and technology, facilitating learning-by-doing, providing imported inputs, and increasing the size of the markets (Grossman and Helpman 1991; Young 1991; Romer 1994).

Numerous empirical studies of the effect of trade liberalization on firm productivity have been undertaken for developing countries in Latin America and Asia.[1] Most of these indicate that the effect is positive.[2] In contrast, and despite the high profile of the topic and the extent of the reforms, very little is known about the effects of trade policy on firm performance in Sub-Saharan Africa. This is a significant information gap. There is evidence in the macro literature that the effects of trade on economic performance vary with the quality of the investment climate and the level of economic development (Chang et al., 2005; Bolaky and Freund, 2006; DeJong and Rippol, 2006). Sub-Saharan Africa is the world's poorest region and the investment climate is generally weak.[3] Hence, the evidence for other regions is unlikely to generalize to Sub-Saharan Africa.

The main reason why there is little research on Africa appears to be lack of suitable data. We have found only two published articles that have examined the

[1] See, for example, Amiti and Konings (2007) on Indonesia; Tybout and Westbrook (1995), Grether (1996), Iscan (1998), and Weiss (1992) on Mexico; Tybout, de Melo, and Corbo (1991), Tybout (1996), and Pavcnik (2002) on Chile; Alan and Morison (2000) on Peru; Ocampo (1994) and Fernandes (2007) on Colombia; Jenkins (1995) for Bolivia; Kim (2000) and Kwak (1994) on Korea; Weiss and Jayanthakumaran (1995) on Sri Lanka; Krishna and Mitra (1998) for a study of India. See Harrison and Rodriguez-Clare (2009) for a survey of the literature.

[2] Notable exceptions include Jenkins (1995) for Bolivia; Ocampo (1994) for Colombia; and Weiss and Jayanthakumaran (1995) for Sri Lanka.

[3] For example, with regards to ease of doing business, the doingbusiness.org website ranks Ethiopia as 107 out of 183 countries.

effect of trade liberalization on manufacturing performance based on firm-level data. The first is that by Harrison (1994), which documents a fall in price-cost margins amongst firms in Cote d'Ivoire following the 1985 trade reform. This is consistent with the hypothesis that openness to imports increases competition in the domestic market. The second is that by Mulaga and Weiss (1996), which, based on data for Malawi 1970-1991, reports mixed results as to the effect of trade reform on productivity growth. Although these studies are interesting and important, the nature of the data limits the range of questions that can be answered. Both studies use data that are very limited in terms of their post-reform coverage, hence the estimated effects might be transitory (see Erdem and Tybout, 2003) or driven by other reforms implemented at the same time as the trade reform. Furthermore, the samples in both studies are fairly small.[4] The Ethiopian government implemented six successive custom tariff reforms between 1993 and 2003 and our data span the period 1997-2005. The staggered nature of the tariff reductions over time and across industrial subsectors enables us to identify the effects of tariffs on firm performance whilst controlling for a wide range of unobservable determinants of productivity.[5] Our estimation sample consists of more than 6,000 observations, which, by the standards of African firm-level datasets, is exceptionally large.[6]

The rest of the paper is organized as follows. The next section describes the context of trade reform in Ethiopia. Section 3 presents the data source and provides descriptive information about the patterns of tariff and import penetration. Section 4 contains the econometric results, and Section 5 provides conclusions.

[4] Harrison (1994) uses data on 246 large and medium sized firms, whereas Mulaga and Weiss (1996) rely on cross-section survey data for 1991-92 that covered only the few surviving large manufacturing establishments that had been in operation since 1973.

[5] No firm-level data are available for the period before the trade reforms were initiated, thus it would not be possible to do a 'before-after' comparison of firm performance along the lines of Harrison (1994).

[6] See Bigsten and Söderbom (2006) for a survey of the literature on manufacturing enterprises in Africa.

2 Context: Trade Reform in Ethiopia

The first Ethiopian development plan in the 1950s advocated an import-substituting strategy based on private ownership. Industrial development gained momentum after the Imperial government introduced measures such as generous tax incentives, high levels of tariff protection, and easy access to domestic credit for domestic production of manufactured goods. The military regime that came into power in 1974 (known as the Derg) nationalized all private large and medium scale manufacturing firms. The regime pursued an import-substituting strategy combined with a command economic system. The industrial development strategy sought to promote industrialization behind high tariff walls and with government ownership. The manufacturing sector was weakened and the private sector was intentionally stifled. The contribution of the private sector to production and employment in medium and large scale manufacturing in 1988/89 was a mere 4 and 8 percent respectively (Central Statistical Agency of Ethiopia, 1990). The public sector also suffered from shortage of raw materials, mainly due to lack of foreign exchange.

After nearly two decades of centralized economic policy a new government took over in 1991, and it has since then undertaken extensive policy reforms to transform the economy into a market oriented one. The government adopted a structural adjustment program in 1992/1993. The reform package was formulated with regard to the complementarity between trade liberalization and macroeconomic management in shaping the reform outcome. The trade reform program aimed at first dismantling quantitative restrictions and then gradually reducing the level and dispersion of tariff rates. Six successive custom tariff reforms were implemented between 1993 and 2003. Table 1 shows the rounds of reforms and tariff rates in detail. In the first round (August 1993) the maximum tariff was reduced from 230 percent to 80 percent. It was then gradually reduced and reached

35 percent in the sixth reform round in 2003. The average weighted tariff rate has been reduced from 41.6 percent to 17.5 percent, and the number of tariff bands has fallen from 23 to 6 including the zero rate band.[7] Other reform measures introduced alongside the trade reforms included foreign exchange market liberalization starting with a massive devaluation in October 1992. Since then the exchange rate has been determined by a weekly auction system. Most price controls and restrictions on private investment have been lifted and a large number of public establishments have been privatized.

3 Data and Descriptive Statistics

For our empirical analysis we match census panel data on manufacturing firms, collected annually by the Central Statistical Agency of Ethiopia, with annual data on tariffs on final goods and imports obtained from the Ethiopian Customs Authority (ECA). We are able to match these datasets for the period 1996/7 (henceforth 1997) to 2004/5 (henceforth 2005). The enterprise census covers every formal manufacturing establishment in the country with 10 and or more workers.[8] We refer to the establishments as firms rather than plants, however this distinction is unlikely to be important as the available data indicate that less than 5% of the firms have more than one branch.[9] There is information in the data on output, inputs (local and imported), sales (local and export), employment, location, own-

[7] The bands are 0; 5; 10; 20; 30; and 35 percent. There are 5,608 tariff lines, out of which 5,424 are subject to ad valorem duties, while the rest are duty free items or prohibited imports. There are no import quotas, but there are import prohibitions on health and environment grounds (MoFED, 2006). Some categories of imported goods are subject to excise tax (3%). Compared to other countries in Sub-Saharan Africa, Ethiopia still has relatively high tariffs. For example, the average tariff level in Sub-Saharan Africa for all goods was 11.8% while for Ethiopia it was 13.0% (World Bank, 2008).

[8] Micro enterprises are thus not represented in the data, and it is unclear how the empirical results reported below generalize to this group of firms.

[9] It should be noted that we only have complete information on the number of branches per firm for 1997 and 1998.

ership type, and a variety of costs. Throughout the analysis, all financial variables are expressed in real terms using unpublished sector-specific deflators supplied to us by the Ministry of finance and Economic Development (MoFED) in Ethiopia. Our estimation sample contains firms from 39 different manufacturing sub-sectors at the 4-digit ISIC level, in 94 locations.

The dataset on tariffs and imports was constructed using unpublished commodity level data made available to us by the ECA. These data were originally organized according to the 6-digit Harmonized System (HS) code. We used concordance information from the World Bank trade website, prepared by Alessandro Nicita and Marcelo Olarreaga, to map the import and tariff information with HS codes at the 6-digit level into 4-digit ISIC product codes, enabling us to match our two datasets.[10] To do this, we define the 4-digit level tariff rate as the unweighted average of the values of duties to imports across the 6-digit level HS commodities within the corresponding 4-digit category.[11] With 39 sectors (at the 4-digit level) observed over 9 years, there is thus plenty of variation in the data in the tariff variable.

Table 2 shows some characteristics of the formal manufacturing sector in Ethiopia, for selected years within our sampling period. Following the reforms the number of firms in the formal manufacturing sector increased by about 41 percent between 1997 and 2004. Employment has also grown but far less than the number of firms. The reduction in mean employment from 137 in 1997 to 105 in 2004, and the big difference between the mean and median employment growth,

[10] There is some ambiguity in the industrial classification in some sectors (mainly machinery and vehicle assembly sectors) forcing us to discard about 200 observations in the manufacturing dataset.

[11] An alternative procedure, sometimes used in the literature, would be to weight the HS 6-digit tariffs by the share of imports. However, as discussed by Rodríguez and Rodrik (2001) and Amiti and Konings (2007), this is potentially problematic since high tariff rates get underweighted because of low import levels. For this reason, we rely on unweighted means. See Amiti and Konings (2007) for further discussion.

indicate that the majority of new entrants were small. In terms of employment, textiles was the leading sector in 1997, followed by food. By 2004, the order had been reversed reflecting a significant contraction in textile sector employment. The leather sector has also seen a reduction in employment.

Figures 1 and 2 show averages and standard deviations of nominal tariffs and import penetration rates, respectively, over the sampling period. These computations are firm-level based, hence tariffs and import penetration rates in sectors with a lot of firms receive a higher weight.[12] Figure 1 confirms that tariffs fell on average, and that the dispersion of tariffs across firms decreased, between 1997 and 2005. The gradual and slow fall in average tariffs is a result of the gradual reform of the tariff regime in Ethiopia, discussed earlier. The reduction in the standard deviation of tariffs (within years) indicates that tariffs became more uniform across sectors as a result of the reforms. We note that uniformity of tariffs has been considered a good policy 'rule of thumb' in the literature on trade reform, although Devarajan et al. (1990) showed that this may not be the case if other taxes are not set optimally. Figure 2 shows that there is no obvious trend in the import penetration rate, defined as the share of imports in the domestic market (domestic output plus imports). This suggests that the tariff reforms have not had much impact on imports, but, as we shall see below, this is actually not the case.

In Table 3 we provide a breakdown by sector, for selected years (for presentational purposes we distinguish industries according to the 3-digit ISIC classification here, however in the econometric analysis below we always define sectors at the 4-digit level). Tariff rates have been declining for all industries except two. In 2005, the industries with the highest tariff rates are garment (34%), footwear (33%)

[12] That is, the computed averages indicate the expected tariff and import penetration rates, for a given year, for a randomly drawn firm from the population of enterprises.

and tobacco (32%). Seven industries have 10 percent or lower tariff rates in the same year. The trend for import penetration differs across industries too, some showing a declining trend and others an increase (e.g. food, garment and wood saw increasing import penetration). Most industries use imported intermediate inputs to some extent, but, somewhat surprisingly given the trade reforms, the dependence on imported inputs has fallen somewhat during the sample period. In 2005 the imported input ratio of total input use was 50 percent and above for 10 out of the 21 industries (3-digit level). Export participation of the Ethiopian manufacturing exhibits an increasing trend, but the majority of the industries have not yet become involved in the export market. The few industries that are significantly involved in the export market are leather, food, footwear and textiles.

We now investigate how tariff rates correlate with several variables of interest. Table 4 shows OLS results based on the following specification

$$X_{jt} = \varphi_1^x T_{jt} + \lambda_t + \epsilon_{jt},$$

where X_{jt} is an outcome variable of interest, T_{jt} is the sector-year tariff rate on final goods, λ_t is a time effect, φ_1^x is a parameter to be estimated, ϵ_{jt} is an error term, and j, t denote sector and time. At this stage we are looking for correlations in the data (conditional on time effects), and all regressions are therefore estimated at the year and sector (4-digit) level. Time dummies are included in all specifications, and standard errors are robust to heteroscedasticity.

We begin by investigating if there is a significant association between tariffs and import penetration ratios. Row (1) in Table 4 shows that the estimated coefficient on tariff is -0.60 and significant at the 1% level. This implies that a reduction in the tariff rate by one percentage point is associated with an increase in the import penetration ratio by 0.60 percentage points. This is consistent with the notion that altering tariffs affects the competition from imports faced by domestic firms. Row (2) shows the results of regressing the Herfindahl index of concentration, defined as the sum of the squared market shares in each four-digit sector, on tariffs and year dummies. The estimated coefficient on the tariff variable is equal to -0.27 and significant at the 5% level. A reduction in tariffs is thus associated with an increase in concentration within sectors. As we shall see below, neither entry rates nor exit rates are significantly correlated with tariffs, hence if tariffs affect market concentration this would suggest that the reduction in tariffs impacts more favorably on large than on small firms (by increasing the share of total output produced by large firms). The analysis in the next section provides some support for this notion.

Next, we consider the relationship between tariffs and labor productivity, defined here as the (sector-year averages of) log of value-added per employee. Row (3) in Table 4 shows that the estimated coefficient on the tariff variable is equal to -1.31 and significant at the 1% level, indicating that a reduction in the tariff rate is associated with an increase in average labor productivity. This is consistent with the idea that trade liberalization leads to productivity gains. The point estimate suggests that a one percentage point reduction in the tariff rate raises value added by about 1.3%. In row (4) we show the results obtained by regressing sector-year averages of the firm-level shares of imported direct inputs in total direct inputs. The estimated coefficient is negative and significant at the 1% level, suggesting that reduced tariffs facilitate for firms to employ imported inputs.

In rows (5) and (6) we show results for regressions in which the average entry and exit rates in the sectors are specified as the dependent variables. We do this in order to check if there is any support for the hypothesis, commonly advanced in the literature on trade liberalization and competitive pressure, that increased competition from imports increases the rate of enterprise turnover. The basic idea of the argument is that competition from imports forces inefficient domestic firms out of business, and frees up resources that get reallocated to new entrants (e.g. see Erdem and Tybout, 2003). The estimated tariff coefficients in the exit and entry regressions are very small and wholly insignificant, however. Thus there is little evidence that tariff rates correlate with enterprise turnover or that the trade liberalization has resulted in a reallocation of resources across firms.

4 Econometric Analysis

4.1 Tariffs and Productivity

Our baseline model is as follows. Firms produce output by means of a Cobb-Douglas production function:

$$y_{ijnt} = \beta^m m_{ijnt} + \beta^k k_{ijnt} + \beta^l l_{ijnt} + \omega_{ijnt} + \eta_{ijnt}, \tag{1}$$

where y_{ijnt} denotes the log of real output, m_{ijnt} is log raw materials, k_{ijnt} is log physical capital, l_{ijnt} is log employment, ω_{ijnt} is total factor productivity (TFP), η_{ijnt} is a measurement error in output, $\beta^m, \beta^k, \beta^l$ are input elasticities, and i, j, n, t denote firm, sector, location and time, respectively. Here we treat β^m, β^k and β^l as constant across firms in all sectors, but we will also consider the results from models in which these coefficients vary freely across sectors. In the special case where the share of raw materials in output is constant, the output production function reduces to a value-added specification. We do report results for value-added specifications below, but focus primarily on the output specification as this is less restrictive. Typically, the results are stronger, in terms of statistical significance, for specifications modeling value-added than for output.

Our main goal is to estimate the effects of trade liberalization on value-added operating through TFP, whilst controlling for other sources of variation in TFP across firms. We thus decompose ω_{ijnt} as follows:

$$\omega_{ijnt} = \theta T_{jt} + \tau_t + \gamma_j + \kappa_n + \alpha_{ijn} + \epsilon_{ijnt}, \tag{2}$$

where τ_t is a time effect common to all firms, γ_j is a sector-level fixed effect, κ_n is a location fixed effect, α_{ijn} is a firm-level fixed effect, and ϵ_{ijnt} is unobserved time varying productivity. Since the tariff variable T_{jt} varies only across sectors and over time, and since we control for both time and firm (or, in some specifications, sector and location) fixed effects, a necessary condition for θ to be identified is that there is variation across sectors in the growth rates of tariff rates.

A common concern in the applied literature on production functions is that the factor inputs are endogenous, i.e. correlated with the error term, in which case the estimated parameters will be biased (Marschak and Andrews, 1944; see Ackerberg, Benkard, Berry and Pakes, 2006 for a recent discussion). A common way to address this is to use proxy variables for unobserved productivity (e.g. Olley and Pakes, 1996; Levinsohn and Petrin, 2003), instrumental variables (e.g. Blundell and Bond, 2000) or some combination of these methods (e.g. Levinsohn and Petrin, 2003; Ackerberg, Caves and Frazer, 2006). Indeed, some recent papers analyzing the effect of trade liberalization on firm-level productivity have used the methods proposed by Olley and Pakes (e.g. Amiti and Konings, 2007; Pavcnik, 2002) and Levinsohn and Petrin (Fernandes, 2007). As pointed out by Ackerberg, Benkard, Berry and Pakes (2006), these methods are not designed to work in the presence of unobserved firm-level fixed effects. At least with our data, assuming there are no fixed effects would be too restrictive, and we therefore eschew these methods.

That the endogeneity of inputs is important to address if the purpose is to obtain consistent estimates of the coefficients on the factor inputs seems clear. It is less likely that non-zero correlation between, say, ϵ_{ijnt} and l_{ijnt} leads to bias in the estimate of θ, our key parameter of interest, if estimated by OLS or fixed effects (the "within" estimator). We obtain our empirical model by plugging (2) into (1):

$$y_{ijnt} = \beta^m m_{ijnt} + \beta^k k_{ijnt} + \beta^l l_{ijnt} + \theta T_{jt} + \tau_t + \gamma_j + \kappa_n + \alpha_{ijn} + \epsilon_{ijnt} + \eta_{ijnt}. \quad (3)$$

Consider first using OLS to estimate this model, with controls for time, sector and location. Consistent estimation of $\beta^m, \beta^k, \beta^l, \theta$ would require materials, capital, labour and tariffs to be uncorrelated with $\left(\alpha_{ijn} + \epsilon_{ijnt} + \eta_{ijnt}\right)$, clearly a strong assumption. The term η_{ijnt}, is assumed to be innocuous measurement error in the dependent variable, thus the potentially troublesome factor is $(\alpha_{ijn} + \epsilon_{ijnt})$. Write the linear projection of $(\alpha_{ijn} + \epsilon_{ijnt})$ on the observed right-hand side variables in (3) as follows:

$$(\alpha_{ijn} + \epsilon_{ijnt}) = \pi_0 + \pi_1 m_{ijnt} + \pi_2 k_{ijnt} + \pi_3 l_{ijnt} + \pi_4 T_{jt} + \tau_t + \gamma_j + \kappa_n + R_{ijnt}$$

where $\pi_0, ...\pi_4$ are coefficients and R_{ijnt} is a zero-mean error term orthogonal to the variables on the right-hand side by definition. The standard worry is that inputs are correlated with the unobserved term $(\alpha_{ijn} + \epsilon_{ijnt})$, in which case π_1, π_2 and π_3 would be different from zero and the OLS estimator applied to (3) would be inconsistent:

$$\begin{aligned} p\lim \hat{\beta}^m_{OLS} &= \beta^m + \pi_1 \\ p\lim \hat{\beta}^k_{OLS} &= \beta^k + \pi_2, \\ p\lim \hat{\beta}^l_{OLS} &= \beta^l + \pi_3. \end{aligned}$$

More importantly from our point of view, if $\pi_4 \neq 0$, the OLS estimate of θ based on (3) would be biased. Why might $\pi_4 \neq 0$? While tariffs are not chosen by the firms, it may be that tariffs are set in response to average performance in the sectors. We control for sector dummies, time effects and factor inputs, but the

OLS estimate of θ would still be biased if, as TFP changes in a sector, the policy maker changes tariffs in response.

One reason sector TFP might change is that the average of the firm fixed effects in the sector changes. This might happen, for example, if firms with low α_{ij} go out of business and get replaced by new firms with higher α_{ij}. If tariffs respond to such a process then there will be endogeneity bias in the OLS estimate. Controlling for firm fixed effects would solve this problem. If shocks to the time varying factor ϵ_{ijnt} are correlated across firms within sectors, and policy makers observe this and change tariffs in response, then within groups and OLS estimates of θ will be biased. In such a situation, an instrumental variable approach might prove useful.

To construct an instrumental variable, we draw on the fact that one explicit objective of the reforms at the outset was to reduce the dispersion of tariff rates across sectors. This is reflected in Figure 1, showing a fall in the standard deviation of tariff rates across sectors over time. Consequently, sectors with high initial tariff rates will have experienced relatively large tariff cuts during the reform period. This, combined with the gradual process by which tariffs were reduced, suggests that initial tariffs interacted with time will be an informative instrument for subsequent tariff levels,

$$T_{jt} = \boldsymbol{X}^1_{ijnt}\boldsymbol{\delta}^1_x + \delta^1_T\left(t \times T_{j0}\right) + e^1_{ijnt}, \tag{4}$$

and that lagged tariffs will be an informative instrument for subsequent growth in tariffs,

$$\Delta T_{jt} = \boldsymbol{X}^2_{ijnt}\boldsymbol{\delta}^2_x + \delta^2_T \times T_{j,t-2} + e^2_{ijnt},$$

where $\boldsymbol{X}^1_{ijnt}$, $\boldsymbol{X}^2_{ijnt}$ are vectors containing all non-instrumented explanatory variables in the main equation, $\boldsymbol{\delta}^1_x$, $\boldsymbol{\delta}^2_x$ are the associated vectors of coefficients, T_{j0} measures the tariff rate in sector j in the initial period, δ^1_T, δ^2_T are scalars (expected to be negative), and e^1_{ijnt}, e^2_{ijnt} are error terms. Our tariff data begin in

1997, thus four years after the reforms were initiated, so we do not have data on T_{j0}. Still, if the residual in the main equation (ϵ_{ijnt}) is serially uncorrelated (a testable assumption), an instrument based on tariffs dated 1997 will be uncorrelated with ϵ_{ijnt} from 1998 and onwards. We therefore estimate fixed effects regressions instrumenting tariffs in the period 1998-2005 with the interaction term between time and the 1997 tariff level. We also estimate the main equation in first differences, in which case we use tariffs lagged two periods as an instrument for ΔT_{jt}.[13] A key assumption is that the policy objective of greater uniformity of tariffs was not meant to favour specific sectors that were expected to have unusually high (or low) subsequent growth rates. A similar identification strategy has been used by Goldberg and Pavcnik (2005) and Amiti and Konings (2007).

How likely is that tariffs are endogenous in this context? Using industry-level data for a sample of six African countries (including Ethiopia), Jones et al. (2008) ask whether the pattern of protection and tariff reform since the early 1990s can be explained by political economy mechanisms, e.g. protection in response to industry lobbies. The evidence suggests that such mechanisms have played a limited role. The authors argue instead that the pattern of tariff reductions was technocratic in structure, noting as we have done above that reductions in average tariffs and in the dispersion of rates were implemented across the board, with larger reductions for higher tariffs. This is consistent with policy reforms being guided by the World Bank, in which case it would seem plausible to argue that policy reforms were essentially exogenous.

[13] Alternatively, we might use the 1997 tariff level as an instrument for all subsequent growth rates starting in 1999. We found this to be a less informative instrument for differences than the tariff level lagged two periods.

4.1.1 Baseline Results

In Table 5 we show our baseline regressions. Columns (1) and (2) show results for the output model while columns (3) and (4) show the results for the value-added specification. All regressions are estimated at the firm-level, and standard errors are firm-level clustered throughout, i.e. robust to heteroscedasticity and serial correlation. The regressions in columns (1) and (3) control for location, industry at the 4-digit level, and common time effects; those in columns (2) and (4) control for firm fixed effects (rendering location and industry dummies redundant) and time effects.

The results for all four specifications indicate a large, negative and statistically highly significant effect of tariffs on productivity. The OLS estimate of the tariff coefficient in column (1) is -0.31, suggesting that a one-percentage point reduction in the tariff rate increases output by 0.3% through higher TFP. The fixed effects estimate is very similar, at -0.32. For reference, we note that Amiti and Konings (2007) report a point estimate equal to -0.21 for Indonesia, based on a similar econometric specification.[14] Taking the coefficient on log raw materials at face value, a tariff coefficient equal to -0.3 in the output specification translates into an effect on value-added by about $1/(1-0.8) \times 0.3 = 1.5\%$.[15] When estimated directly, the value-added models indicate a similar order of magnitude for the effect of tariffs on value-added. The t-values on the tariff coefficients are considerably higher in the value-added regressions than in the output models, possibly because

[14] Amiti and Konings (2007) show results both for a two-step procedure, in which TFP is calculated as the residual in the production function estimated separately, and a one-step procedure in which log output is regressed on the tariff variable controlling for labour, capital, materials and other factors. The latter approach is thus the same as ours. The two methods produce virtually the same estimate (-0.21), see Table 4 col.1 and Table 5 col.4 in Amiti and Konings (2007).

[15] The sample median of the share of raw materials in output is actually about 0.6. This suggests the OLS estimate of the coefficient on raw materials may be upward biased, presumably because materials is a highly endogenous variable.

the value-added specifications are more restrictive. Due to likely endogeneity bias we do not interpret the coefficients on materials, capital and labor as estimates of the 'structural' technology parameters in the production function. However we note that the results are broadly in line with those reported in previous studies of manufacturing firms in Africa (Söderbom and Teal, 2004; Baptist and Teal, 2008).

4.1.2 Productivity and Tariffs: Robustness to Endogeneity

Next we investigate if our baseline results are robust to treating the tariff variable as endogenous. Two-stage least squares estimates are shown in Table 6. Columns (1)-(4) present results for models estimated in first differences, while columns 5-6 show levels results (with controls for sector and location fixed effects, and firm fixed effects, respectively). All regressions control for time dummies, and standard errors are clustered at the firm-level.

The specification in column (1) treats factor inputs as econometrically exogenous, which is a strong assumption but provides us with a useful benchmark. Tariffs differenced is instrumented with tariffs lagged two periods. The estimated tariff coefficient is equal to -0.59 and significant at the 10% level, but not at the 5% level. The point estimate is higher than those reported in Table 4, in which the tariff variable is assumed exogenous. The first-stage regression (not shown) indicates that lagged tariffs is a strong instrument for current differences, and we can reject the null hypothesis of underidentification at the 1% level. Based on a Hausman type test for exogeneity we do not reject the hypothesis that tariffs are exogenous. There is strong evidence that the differenced residual follows a MA(1) process, but there is no evidence of second order serial correlation. This suggests that the levels residual is serially uncorrelated, a key assumption for our present identification strategy.

As noted above, the assumption that the factor inputs (capital, labour, materials) are exogenous is potentially strong. For example, if tariffs affect productivity and inputs respond to changes in productivity, lagged values of tariffs will generally not be valid instruments if tariffs are endogenous to unobserved productivity. To allow for a situation in which inputs are endogenous to productivity and tariffs are a determinant of productivity while at the same time endogenous to unobserved productivity, we must instrument the inputs as well as the tariff variable. This, however, has proven difficult with the data at hand. We began by using values of all the inputs lagged two periods as instruments for current differences (the economic argument underlying this strategy is that inputs change slowly, perhaps because of adjustment costs or slow changing factor prices; see Bond and Söderbom, 2005). However, this procedure gave very imprecise estimates of the parameters in the main equation, and we could not reject the hypothesis that this model is underidentified. Our instruments for the three inputs and the tariff variable (differenced) are clearly too weak to be useful.

In view of these difficulties, we decided to re-estimate the model with constant returns to scale imposed, on the grounds that this restriction results in one less endogenous explanatory variable, and that constant returns to scale is a common result for African firms (e.g. Söderbom and Teal, 2004; Baptist and Teal, 2008). Results are shown in Table 6, column (2). Imposing constant returns to scale helps in the sense that we can now firmly reject the null hypothesis that the model is underidentified. As expected, all t-values are much lower than in column (1). The estimated coefficient on the tariff variable is equal to -0.54, thus similar to results in previous specifications, but with a t-value equal to 1.35 we cannot reject the hypothesis that the coefficient is zero. Again, however, we cannot reject the hypothesis that tariffs are exogenous, suggesting that we are instrumenting when in fact we do not have to. In column (3) we switch from using inputs dated

$t-2$ to $t-1$ as instruments, which amounts to assuming that input levels are predetermined. We continue to treat tariffs as endogenous (i.e. we use tariffs dated $t-2$ as an instrument). This improves the precision with which all coefficients of interest are estimated, and the tariff coefficient is now negative and significant at the 5% level. We do not reject the hypothesis that tariffs are exogenous. There is evidence of first-order, but not higher order serial correlation, in the differenced residuals, and the validity of the overidentifying restriction appears marginal.

Columns (4)-(5) show results for levels regressions in which tariffs are instrumented as indicated in equation (4), with year 1997 corresponding to $t = 0$. The within transformation implies that the instrument must predate the sampling period. Under the assumption that ϵ_{ijnt} is serially uncorrelated, $(t \times T_{j0})$ would be a valid instrument for tariffs provided the estimation sample period is $t = 1, 2, ...S$.[16] The estimates of the tariff coefficient are similar to the first difference results. With the specification used in column (4), in which we control for sector and location effects, the tariff coefficient is significant at the 10% level, while in that shown in column (5), which controls for firm fixed effects, the effect is close to significant at the 10% level. In neither model do we reject the hypothesis that tariffs are exogenous.

4.1.3 Market Concentration, Technology and Trade Patterns

The main conclusion from the analysis in the previous sub-section is that there is no strong evidence that the fixed effects estimates of the tariff effect are biased by endogeneity. Based on what we know about the politics of the trade liberalization in Ethiopia, this seems like a reasonable result. We now extend the baseline

[16] Unfortunately it is not straightforward to test for serial correlation in the levels residual based on fixed effects estimates if the panel is unbalanced. The within transformation implies the estimated residual will exhibit serial correlation under the null, that depends on the length of the panel (Wooldridge, 2002, p.275). Thus, with an unbalanced panel it is not clear what the null hypothesis should be.

specification in various ways. First we relax the restriction that the coefficients on the factor inputs are constant across sectors. Heterogeneity in technology across sectors is likely to result in differing coefficients on materials, capital and labor. With 39 sectors included in the sample, the output regressions in Table 5 may have as many as 114 omitted interaction terms.[17] From our point of view, whether this is a serious problem depends on whether these omissions result in bias in the estimated tariff effect. There will be such bias if the parameter π_4 is different from zero in the following linear projection:

$$u_{ijnt} = \pi_0 + \pi_1 m_{ijnt} + \pi_2 k_{ijnt} + \pi_3 l_{ijnt} + \pi_4 T_{jt} + \tau_t + \gamma_j + \kappa_n + R_{ijnt}$$

where

$$u_{ijnt} = \alpha_{ijn} + \epsilon_{ijnt} + \left(\beta_j^m - \bar{\beta}^m\right) m_{ijnt} + \left(\beta_j^k - \bar{\beta}^k\right) k_{ijnt} + \left(\beta_j^l - \bar{\beta}^l\right) l_{ijnt}.$$

This could be the case, for example, if tariffs are set low in sectors with high returns to scale (such sectors will tend to have high values of u_{ijnt}). Column 1 in Table 7 summarizes fixed effects results based on a model that includes a full set of industry-materials, industry-capital and industry-labor interaction terms. We are not particularly interested in the sector specific estimates of the input coefficients (these are possibly biased by endogeneity anyway, as already discussed), so we report averages of these estimates along with t-values associated with the null hypothesis that these averages are equal to zero. The key finding is that the tariff coefficient remains negative and significantly different from zero at the 5% level. Thus, allowing for industry heterogeneity in the input coefficients has a very small effect on the tariff coefficient.

A common concern in the literature is that it is hard to estimate the effects of tariff reforms on productivity and mark-ups separately (e.g. Harrison, 1994).

[17] That is, 39-1 sector dummies interacted with materials, capital and labor.

In particular, changes in mark-ups resulting from the trade liberalization may show up as productivity effects unless properly controlled for. To address this, we add the Herfindahl index of industry concentration with respect to sales from domestic firms to the set of explanatory variables. We return to the baseline specification with common input coefficients across sectors, since we have found that ignoring industry heterogeneity in this context does not lead to bias. Results are shown in Table 7, column (2). The Herfindahl index, measured at the 3-digit industry level so as to ensure that a reasonable number of observations underlies the computations, comes in with a negative and significant coefficient, indicating that an increase in industry concentration reduces firm-level productivity, other factors held constant. A similar result is reported for Indonesia by Amiti and Konings (2007). Importantly, the coefficient on tariffs remains negative and highly significant. If we add to this specification an interaction term between the tariff variable and the Herfindahl index, the associated coefficient is wholly insignificant (t-value 0.74; results not shown). Thus, there is no evidence that the effects of tariff cuts on productivity depend on industry concentration. Our main conclusion from this analysis is that changes in the mark-up cannot account for the effect of tariffs on productivity that we observe.

Trade liberalization is sometimes accompanied by a depreciation of the real exchange rate[18], which may lead to increased demand and productivity gains for exporting firms. For Ethiopia, few firms export and, over our sampling period, the real exchange rate is relatively stable.[19] Still, it is of interest in and of itself to investigate if changes in the real exchange rate have impacted firm-level productivity in the country. In column 3 we add the real exchange rate to the baseline specification (which rules out including time dummies in the regression). The coefficient

[18] See for example Fernandes (2007) for an analysis of Colombia.

[19] In the early 1990s, however, there was a massive devaluation, as noted above.

on the real exchange rate is positive and weakly significant (at the 10% level), suggesting that a real depreciation tends to raise productivity.[20] As expected, the tariff coefficient does not change very much, since the time dummies would have picked up macro effects in the previous models. We have also investigated if the tariff effect varies with real exchange rate, by adding to the baseline model an interaction term between the tariff variable and the real exchange rate. The associated coefficient is insignificant however (t-value 1.31; results not shown). Hence, while there is some evidence that movements in the real exchange rate have impacted productivity, there is no evidence that the effects of lower tariffs estimated above are due to changes in the real exchange rate.

As discussed in the introduction, the most commonly advanced explanation in the literature as to why reducing output tariffs produces productivity gains is that import competition becomes tougher. However, as noted by Amiti and Konings (2007), an alternative channel may be through cheaper or better imported inputs. To investigate this, Amiti and Konings construct measures of the input tariffs faced by Indonesian firms, and find that the effect of reduced input tariffs on productivity is typically larger than that of reduced output tariffs. These authors also find that firms that import their inputs enjoy a larger productivity gain from trade liberalization than nonimporting firms, which is consistent with the notion that benefits accrue primarily on the input side. We do not have data on input tariffs and therefore cannot document the effects of input tariffs on productivity directly. However, we do have information about the share of imported inputs in total inputs. If the trade liberalization impacts performance primarily through better and/or cheaper inputs, and if (as seems likely) output tariffs and input tariffs are positively correlated, we would expect to find a larger tariff coefficient

[20] Naturally, this result could be driven by other macro variables that we cannot control for.

for firms with large shares of imported inputs. We therefore add to the baseline specification the share of imported inputs on its own and interacted with the tariff variable. Results are shown in column (4). The coefficient on the imported inputs interacted with tariffs is positive (hence the 'wrong' sign if the prior is that the tariff effect is stronger for firms with a lot of imported inputs) and wholly insignificant. Of course, without data on input tariffs we cannot rule out the possibility that the estimates of the tariff effect at least partly reflect mechanisms operating through better or cheaper inputs. Still, the results in column (4) are at least suggestive that such mechanisms are not very strong.

In a similar vein, we explore whether the tariff effect varies depending on exporting status. Several studies have documented a positive relationship between exporting and productivity for African firms (Bigsten et al., 2004; Van Biesebroeck, 2005; Bigsten and Gebreeyesus, 2009). It would seem likely that trade liberalization induces more exporting. Could it be, then, that the significant tariff effect in our baseline specification is an artefact of not controlling for exporting? The results, shown column 5 in Table 7, indicate no strong effect of exporting on productivity, neither on its own or in conjunction with lower tariffs. The estimated tariff coefficient, interpretable as measuring the average effect of tariffs on productivity, is negative and significant at the 1% level.[21]

4.1.4 Heterogeneous Productivity Gains: Firm Size and Non-Linearities

Table 8 shows results for models in which we allow for heterogeneity in the tariff effect depending on firm size as well as nonlinear tariff effects. All regressions control for firm fixed effects. To investigate if the tariff effect varies with firm size, we define firms as small (large) if employment is lower (higher) than the

[21] We subtract the sample mean from the exports dummy in order to facilitate interpretation of the tariff coefficient (i.e. the exports dummy entering the regression has zero mean).

sample median and interact these size dummies with the tariff variable. Results, shown in column 1, suggest a stronger effect of tariffs on productivity amongst the large firms. For this group, the estimated tariff coefficient is -0.40 and significant at the 1% level, whereas for the small firms the tariff effect is only half as large and insignificant at conventional levels. This is qualitatively similar to findings reported for Colombia by Fernandes (2007). However, based on a t-test we do not reject the null hypothesis that the tariff coefficient is common for the two size groups (the p-value is 0.13). We thus interpret the results as providing some, but not very strong, support for the idea that large firms have benefited more than small firms from the trade reforms.

In column 2 we add tariffs squared to the baseline model in order to test for nonlinearities. The estimated coefficient on tariffs is positive while that on the squared term is negative, indicating an inverse-u type relationship between tariffs and productivity. Both coefficients are significantly different from zero at the 1% level. These results suggest that high tariffs are particularly distortionary. For example, evaluated at a tariff rate of 40%, the partial effect of a small change in tariffs on productivity is -1.13. Figure 3 shows how predicted productivity varies with tariffs based on this model (Quadratic). In column 3 we interact tariffs and tariffs squared with the size dummies to test for differences across firm size. The estimated coefficients on tariffs and tariffs squared are very similar for the two size groups and we don't reject the null hypothesis that the tariff-productivity relationship is the same for the two groups.

In columns (4) and (5) we show results from specifications in which we model the relationship between tariffs and productivity using a piecewise linear spline function:

$$\omega_{ijnt} = \theta_1 T_{jt} + \sum_{k=2}^{K} \theta_k \max\left(T_{jt} - c_k, 0\right) + \tau_t + \gamma_j + \kappa_n + \alpha_{ijn} + \epsilon_{ijnt},$$

where K denotes the number of nodes, $c_k > 0$ determines the position of the kth node, and $\theta_1, \theta_2, ..., \theta_K$ are parameters determining the productivity tariff profile. The coefficient θ_1 is interpretable as the slope of the profile in the range $0 \leq T \leq c_1$, while θ_k, $k = 2, 3, ..., K$, is interpretable as the change in the slope of the profile that results from moving from the interval $\{c_{k-2}, c_{k-1}\}$ to $\{c_{k-1}, c_k\}$, where $c_0 = 0$. The slope of the productivity function in the interval $\{c_{L-1}, c_L\}$ is thus given by $\theta_1 + \sum_{k=2}^{L} \theta_k$. Hence, if $\theta_2 = \theta_3 = ... = \theta_K = 0$ the productivity function is linear.

Column (4) shows results for a specification in which there are nodes at 0.05, 0.10,...,0.55 (the highest tariff recorded in the data is 0.54). Several of the coefficients $\theta_2, ..., \theta_K$ are significantly different from zero, indicating a nonlinear relationship. To facilitate interpretation we show in Figure 3 how predicted productivity varies with tariffs based on these results (Spline I). For low and intermediate tariff levels the productivity-tariff curve is wiggly and there is no obvious systematic pattern, but at high tariff levels there is a strong negative relationship between tariffs and productivity. In column (5) we show results for a less flexible specification allowing for a single kink in the function at the point where the tariff rate is 30%. These results are easier to interpret and less noisy. The estimated coefficient on tariffs below 30% is equal to -0.027 but wholly insignificant, while at tariffs above 30% the estimated partial effect is -1.14 and significant at the 1% level. The results are illustrated in Figure 3 (Spline II). For about 20% of the observations in our sample tariffs exceed 30%, hence, based on the results in column (5), the sample average of the tariff effect is $0.8 \times (-0.027) + 0.2 \times (-0.027 - 1.108) = -0.25$ (significant at the 1% level). This is similar to the estimated tariff coefficient in the linear specification reported in Table 5, column 2 (-0.32). By adopting a nonlinear specification we thus learn that most of this effect is driven by the strong negative impact of tariffs on productivity at high tariff levels.

4.2 Input Decisions

So far we have focused on the effects of the tariff cuts on firm-level productivity. It is perfectly possible, however, that the trade liberalization has other important effects on firm performance too. For example, one might hypothesize that firms grow as a result of the trade liberalization, in response to higher marginal products for the inputs. Table 9 shows results from fixed effects regressions modelling employment, the capital-labour ratio (both in logs) and the share of imported inputs as dependent on tariffs. Column (1) shows that there is some evidence (at the 10% significance level) that the reduction in tariffs has resulted in a *reduction* in employment. The result that firms are becoming smaller and more productive at the same time is consistent with the hypothesis that the trade liberalization has increased competition in the domestic market. For example, if the market size is fixed, the entry of new competitors is likely to lead to smaller market shares for the incumbents, thus disincentivizing their growth. This idea is consistent with the negative relationship between import penetration ratios and tariffs observed in the data (Table 4). It is also consistent with simulations based on a computable industrial evolution model, reported by Erdem and Tybout (2003).

The results in column (2) indicate that the tariff cuts have led to higher capital-labour ratios (significant at the 5% level). This can be interpreted as indicating that tariff reductions have led to a lower cost of capital relative to labour, perhaps because imported equipment has become cheaper. However, the results in column (3) indicate no significant effect of tariffs on the share of imported inputs, and so possibly the effect of the tariff cuts on the relative prices faced by manufacturing firms has not been very strong.[22] An alternative explanation as to why firms

[22] Recall from Table 4 that there is a negative and highly significant correlation between tariffs and the share of imported inputs across sectors. This result disappears once we control for sector or firm fixed effects. We conclude that the significant negative correlation between these

appear to have become more capital intensive in response to lower tariffs is that this is a short-term effect: capital is a less flexible input than labour, and the downward adjustment of capital takes longer than that for labour.

5 Conclusions

In this study we use firm-level data for the Ethiopian manufacturing sector matched with commodity-level data on tariffs to examine the effect of trade liberalization on firm performance. The period spanned by our data, 1997-2005, was one of successive tariff reforms. These were accompanied by other broad reform measures such as privatization, deregulation of prices, and liberalization of exchange rate and financial markets. However, unlike these broader reform measures the pace of tariff reforms varied across sectors and over time, enabling us to estimate the effects of tariff cuts on firm performance whilst controlling for macroeconomic effects that can be thought of as approximately constant across sectors.

Using a three-factor output Cobb-Douglas production function as our baseline model, we find relatively large positive effects of tariff reductions on total factor productivity. This finding is robust to treating tariffs as an endogenous variable, and to various generalizations of the baseline model (e.g. models allowing for industry-specific input coefficients; industry concentration; the extent of imported inputs; and the extent of exporting). Results based on a two factor value-added specification (which is more restrictive than the output model) are even stronger in terms of their statistical significance. There is some evidence that the tariff effect is stronger for large than for small firms, although when tested formally we cannot reject the hypothesis that the effect is common across the two size groups. It is clear, however, that the negative relationship between tariffs and productivity

variables in the cross-section is driven by heterogeneity across firms/sectors in the underlying intensity with which imported inputs are used.

is primarily driven by mechanisms operating at high tariff levels. Models allowing for nonlinear tariff effects strongly indicate the presence of nonlinearities. The results from a specification containing a simple piecewise linear spline function with a single kink at a tariff level of 30% indicate that the tariff effect below 30% is close to zero, while at tariffs above 30% the estimated partial effect is -1.14 and significant at the 1% level.

We also investigate that relationship between tariffs and employment, the capital-labour ratio and the share of inputs imported. The results suggest that lower tariffs are associated with a fall in average firm size. This is consistent with the notion that the trade liberalization has increased competition in the domestic market, resulting in higher productivity and lower market shares for domestic firms. In terms of firm profitability, these two effects may well offset each other. Indeed, assuming that entry and exit decisions reflect profitability shocks, we can interpret the fact that there is no significant effect of tariffs on domestic entry or exit rates as suggesting that the trade liberalization has not affected the average profitability of firms.

For about 20% of the observations in our sample, and 11% in the final year of the panel, tariffs exceed 30%. Our results indicate that such firms will have benefitted from the trade liberalization, and would benefit from further cuts in cases where tariffs remain high. The very low effect estimated for tariffs less than 30% suggests that low and intermediate tariffs may not be overly distortionary, however. For a long time it has been part of the folk wisdom of development economics that high tariffs reduce consumer welfare and harm productive efficiency. Our study indicates that the latter effect is modest for moderate tariffs but very strong for excessive tariffs..

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Table 1: Tariff reform steps in Ethiopia (1993-2003)

Rounds of reforms	Year	Maximum tariff	Average tariff	Number of tariff bands
Before reform	Before 1993	230	41.6	23
1st round	August 1993	80		
2nd round	January 1996	60		
3rd round	_______ 1997	60		
4th round	January 1998	50	21.5	
5th round	December 1998	40	19.5	
6th round	January 2003	35	17.5	6

Source: MoFED (Study on Ethiopia's Industrial sector Effective Rate of Protection, December, 2006 – MoFED mimeo)

Table 2: Number of establishments and employment in Ethiopia's manufacturing sector

Sector	Firms			Employment								
	Number of firms		Growth	Total Employment		Growth	Sector share		Average firm size		Median firm size	
	1997	2004	1997-04	1997	2004	1997-04	1997	2004	1997	2004	1997	2004
Food	179	294	64%	26,926	31,238	16%	28%	30%	150	106	21	24.5
Textile	59	73	24%	31,839	26,677	-16%	33%	25%	540	365	51	58
Leather	61	62	2%	8,226	7,575	-8%	9%	7%	135	122	27	49.5
Wood	132	185	40%	5,680	6,822	20%	6%	7%	43	37	20.5	16
Paper	46	73	59%	5,122	6,929	35%	5%	7%	111	95	24.5	35
chemical	64	87	36%	6,124	9,306	52%	6%	9%	96	107	36	59
Non-metallic	89	119	34%	6,745	9,170	36%	7%	9%	76	77	17	19
Fabricated metal	72	103	43%	4,377	6,594	51%	5%	6%	61	64	20.5	30
Total	703	997	42%	95,992	105,095	10%			137	105	23	26

Note: The figures are based on the Ethiopian census data.

Table 3: Tariff, import penetration, imported inputs and export ratios

	Average Tariff (% of CIF import value)			Import penetration ratio			Average share of imported inputs in total inputs			Average share of exports in sales		
	1997	2001	2005	1997	2001	2005	1997	2001	2005	1997	2001	2005
Food	29	29	24	0.08	0.20	0.20	0.10	0.14	0.09	0.01	0.09	0.12
Other food	30	17	22	0.22	0.38	0.21	0.06	0.10	0.03	0.28	0.08	0.26
Beverage	18	12	10	0.03	0.05	0.02	0.43	0.36	0.52	0	0	0
Tobacco	na	26	32	na	0.03	0.06	0.84	0.73	0.43	0	0	0
Textiles	27	25	16	0.30	0.22	0.20	0.42	0.38	0.50	0.04	0.11	0.08
Garment	46	39	34	0.40	0.55	0.86	0.31	0.07	0.21	0.04	0.02	0.00
Leather	29	31	29	0.08	0.04	0.03	0.13	0.12	0.11	-	0.67	0.76
Footwear	48	39	33	0.31	0.23	0.07	0.51	0.51	0.45	0	0.01	0.12
Wood	7.3	8.3	3.2	0.52	0.60	0.71	0.45	0.63	0.48	0	0	0
Furniture	19	20	26	0.29	0.24	0.19	0.22	0.35	0.42	0	0	0
Paper	12	12	9.2	0.50	0.47	0.28	0.90	0.91	0.96	0	0	0
Printing	12	8.7	9.8	0.44	0.35	0.14	0.48	0.37	0.74	0	0	0
Ind. chemicals	6.8	3.8	3.3	0.92	0.94	0.90	0.32	0.47	0.52	0	0	0
Other chemicals	20	15	9.7	0.55	0.57	0.52	0.79	0.77	0.86	0	0	0
Rubber	14	10	12	0.64	0.54	0.53	0.98	0.98	1.00	0	0	0
Plastic	30	27	22	0.39	0.32	0.32	0.92	0.95	0.97	0	0	0
Glass	18	17	11	0.73	0.70	0.75	0.14	0.34	0.22	0	0	0
Non-metal	12	17	21	0.15	0.07	0.05	0.06	0.08	0.15	0	0	0
Basic iron	6.8	6.9	7.6	0.70	0.54	0.43	0.99	0.99	0.61	0	0	0
Fabricated metal	15	11	12	0.77	0.99	0.49	0.77	0.77	0.81	0	0	0

Table 4: Tariffs and Outcome Variables of Interest

		Tariff		
	Dependent variable	Coefficient	t-value	R-squared
(1)	Import penetration ratio	-0.598	(4.20)**	0.07
(2)	Herfindahl index	-0.267	(2.28)*	0.02
(3)	log Value-added per employee	-1.313	(2.98)**	0.03
(4)	Share of imported inputs	-0.391	(2.76)**	0.02
(5)	Entry rate	-0.081	(1.09)	0.10
(6)	Exit rate	0.071	(0.94)	0.10

Note: The estimation method is OLS. The regressions are estimated at the (4-digit) sector-year level. Year dummies and a constant are included in all regressions. t-values are based on standard errors robust to heteroskedsticity. * significant at 5% level; ** significant at 1% level. The number of observations is 342 in all specifications except in (6), for which it is 303 (data on exit for the last wave of the panel are not available).

Table 5: Tariffs and Firm-Level Productivity: Baseline specifications

	(1) log Output	(2) log Output	(3) log Value-added	(4) log Value-added
log Raw Materials	0.827	0.791		
	(113.62)**	(53.70)**		
log Physical Capital	0.032	0.069	0.199	0.199
	(7.05)**	(4.52)**	(12.40)**	(4.34)**
log Employment	0.173	0.175	1.032	0.851
	(15.71)**	(8.47)**	(36.67)**	(13.93)**
Tariff (4-digit level)	-0.305	-0.317	-1.499	-1.702
	(2.84)**	(2.74)**	(4.11)**	(4.45)**
Time effects	Yes	Yes	Yes	Yes
Location effects	Yes	Redundant	Yes	Redundant
Firm fixed effects	No	Yes	No	Yes
Industry effects (4-digit level)	Yes	Redundant	Yes	Redundant
Observations	6096	6096	6096	6096
Firms	1,700	1,700	1,700	1,700
R-squared	0.98	0.81	0.80	0.16

Note: t-statistics based on robust standard errors clustered at the firm-level in parentheses. *significant at 5% level; ** significant at 1% level. Overall R-squared reported for models without firm fixed effects; within R-squared reported for models with firm fixed effects.

Table 6: Tariffs and Firm-Level Productivity: Two-Stage Least Squares Results

	First Differences			Levels (location & sector effects)	Within (firm fixed effects)
	(1)	(2)	(3)	(4)	(5)
log Raw Materials	0.757	0.821	0.710	0.833	0.800
	(46.33)**	(8.39)**	(22.71)**	(119.7)**	(51.85)**
log Physical Capital	0.036	0.036	0.114	0.030	0.051
	(1.67)	(0.62)	(2.98)**	(6.47)**	(2.75)**
log Employment	0.127	0.143	0.176	0.171	0.178
	(6.05)**	(1.38)	(4.49)	(16.10)**	(8.32)**
Tariff	-0.590	-0.543	-0.757	-0.536	-0.688
	(1.71)	(1.35)	(2.06)*	(1.77)$^{+}$	(1.51)
Tests (p-values)					
Tariff exogenous	0.28	0.42	0.20	0.14	0.21
Underidentification	0.00	0.00	0.00	0.00	0.00
Overid restrictions		0.17	0.08		
Constant returns	0.01			0.00	0.18
Autocorrelation, m1	0.00	0.00	0.00		
Autocorrelation, m2	0.64	0.64	0.65		
Inputs	Exogenous	Endogenous	Predetermined	Exogenous	Exogenous
CRS imposed	No	Yes	Yes	No	No
Exclusion restrictions	Tariff(t-2)	Tariff(t-2) Inputs(t-2)	Tariff(t-2) Inputs(t-1)	Initial tariff x time	Initial tariff x time
Observations	3031	3031	3031	5495	5495

Note: t-statistics based on robust standard errors clustered at the firm-level in parentheses. * significant at 5% level; ** significant at 1% level. The specification in (5) controls for location and industry effects. Time dummies are included in all specifications. m1 and m2 are the Arellano-Bond (1991) tests for first and second order serial correlation in the differenced residuals. The number of observations varies because of how the instruments are defined.

Table 7: The Effects of Market Concentration, the Real Exchange Rate and the Firm's Trade Patterns on Productivity

	(1)	(2)	(3)	(4)	(5)
log Raw Materials	0.777	0.791	0.798	0.801	0.791
	(54.22)**	(53.67)**	(55.59)**	(65.51)**	(54.07)**
log Physical Capital	0.045	0.068	0.065	0.065	0.068
	(2.66)**	(4.49)**	(4.41)**	(4.39)**	(4.50)**
log Employment	0.170	0.174	0.177	0.163	0.175
	(8.85)**	(8.47)**	(8.70)**	(8.97)**	(8.49)**
Tariff (4-digit level)	-0.240	-0.349	-0.282	-0.290	-0.322
	(1.98)*	(3.01)**	(2.47)*	(2.40)*	(2.74)**
Herfindahl index (3-digit level)		-0.481			
		(2.50)*			
Real exchange rate (2000 = 1.00)			0.086		
			(1.84)+		
Tariff x Share imported inputs				0.127	
				(0.53)	
Share imported inputs				0.003	
				(0.07)	
Tariff x Any exports					0.120
					(0.479
Any exports					-0.025
					(0.30)
Time effects	Yes	Yes	No	Yes	Yes
Firm effects	Yes	Yes	Yes	Yes	Yes
Industry heterogeneity in input coefficients (4 digit level)	Yes	No	No	No	No
Observations	6,096	6,096	6,096	6,088	6,096
Firms	1,700	1,700	1,700	1,700	1,700
R-squared	0.82	0.81	0.81	0.81	0.81

Note: The dependent variable is log output. t-statistics based on robust standard errors clustered at the firm-level in parentheses. * significant at 5% level; ** significant at 1% level. Imported inputs and exports are centered on their sample means, so as to facilitate interpretation of the tariff coefficient.

Table 8: Heterogeneous Productivity Gains: Firm Size and Nonlinear Effects

	(1)	(2)	(3)	(4)	(5)
log Raw Materials	0.791	0.789	0.789	0.786	0.789
	(53.80)**	(53.66)**	(53.70)**	(52.91)**	(53.66)**
log Physical Capital	0.068	0.071	0.069	0.068	0.071
	(4.46)**	(4.67)**	(4.59)**	(4.49)**	(4.71)**
log Employment	0.183	0.176	0.184	0.179	0.176
	(7.74)**	(8.50)**		(8.69)**	(8.46)**
Tariff		0.920		0.290	-0.027
		(2.70)**		(0.11)	(0.19)
Tariff x Small	-0.194		0.948		
	(1.39)		(1.74)		
Tariff x Large	-0.400		0.792		
	(3.09)**		(1.81)		
Tariff squared		-2.567			
		(3.77)**			
Tariff squared x Small			-2.380		
			(2.25)*		
Tariff squared x Large			-2.509		
			(2.99)**		
max(tariff-.05,0)				1.589	
				(0.56)	
max(tariff-.10,0)				-3.422	
				(2.99)**	
max(tariff-.15,0)				2.908	
				(2.89)**	
max(tariff-.20,0)				-2.594	
				(2.41)*	
max(tariff-.25,0)				1.266	
				(1.07)	
max(tariff-.30,0)				1.149	-1.108
				(0.96)	(3.43)**
max(tariff-.35,0)				-3.742	
				(2.44)*	
max(tariff-.40,0)				0.669	
				(0.44)	
max(tariff-.45,0)				2.133	
				(0.74)	
max(tariff-.50,0)				-6.022	
				(1.22)	
Large (emp > 26)	0.022		0.019		
	(0.62)		(0.30)		

The table continues on the next page.

Table 8 Continued

	(1)	(2)	(3)	(4)	(5)
H_0: Quadratic, Small (p-value)			0.022		
H_0: Quadratic, Large (p-value)			0.000		
H_0: Quadratic common across size (p-value)			0.257		
Time effects	Yes	Yes	Yes	Yes	Yes
Firm effects	Yes	Yes	Yes	Yes	Yes
Observations	6,096	6,096	6,096	6,096	6,096
Firms	1,700	1,700	1,700	1,700	1,700
R-squared	0.81	0.81	0.81	0.81	0.81

Note: The dependent variable is log output. t-statistics based on robust standard errors clustered at the firm-level in parentheses. * significant at 5% level; ** significant at 1% level.

Table 9: Tariffs and Input Decisions

	(1) Log Employment	(2) Log Capital-Labor Ratio	(3) Share of imported inputs
Tariff	0.306	-0.504	0.006
	(1.74)+	(2.14)*	(0.09)
Time effects	Yes	Yes	Yes
Firm effects	Yes	Yes	Yes
Observations	6,096	6,096	6,088
Firms	1,700	1,700	1,700
R-squared	0.02	0.01	0.005

Note: t-statistics based on robust standard errors clustered at the firm-level in parentheses. * significant at 5% level; ** significant at 1% level.

Figure 1: Trends in the Mean and Standard Deviation of Tariffs

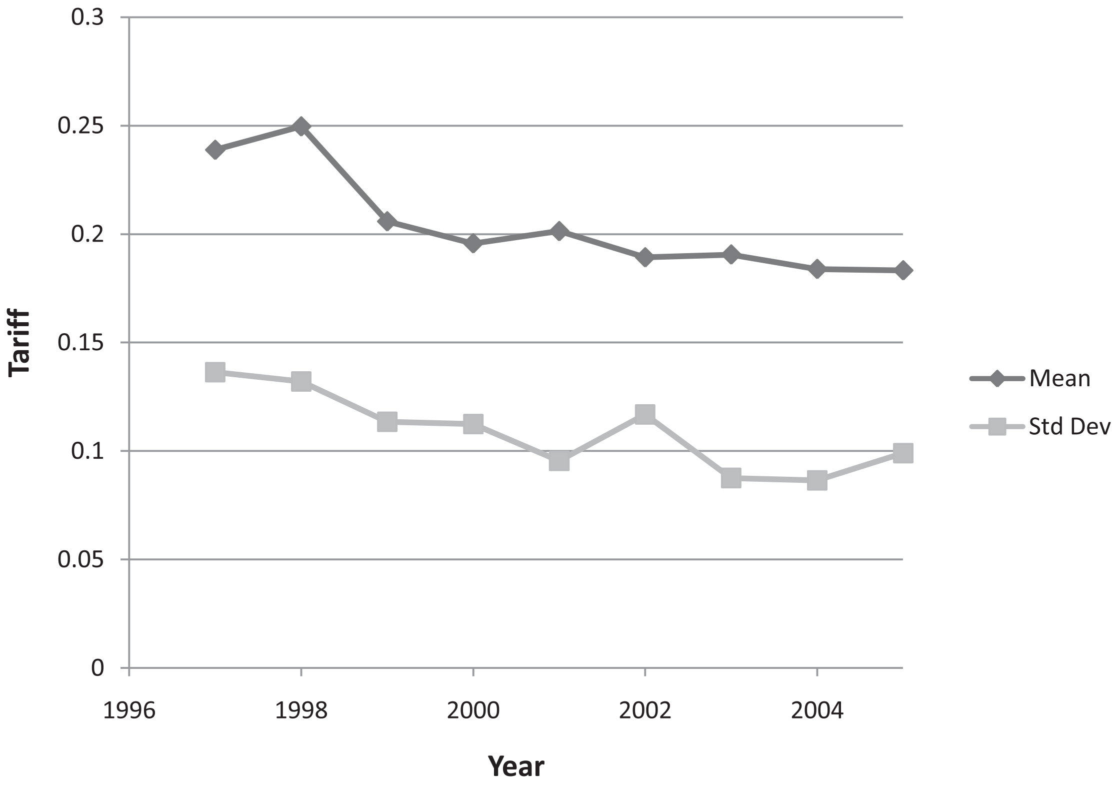


Figure 2: Trends in the Mean and Standard Deviation of Import Penetration Ratios

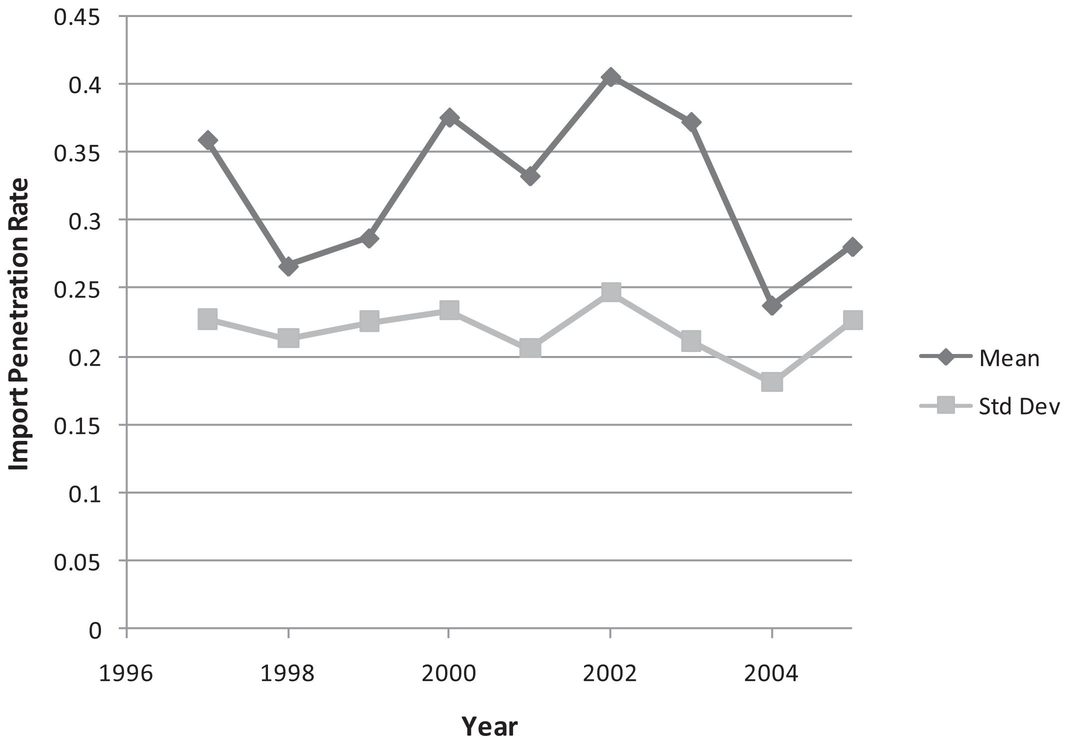
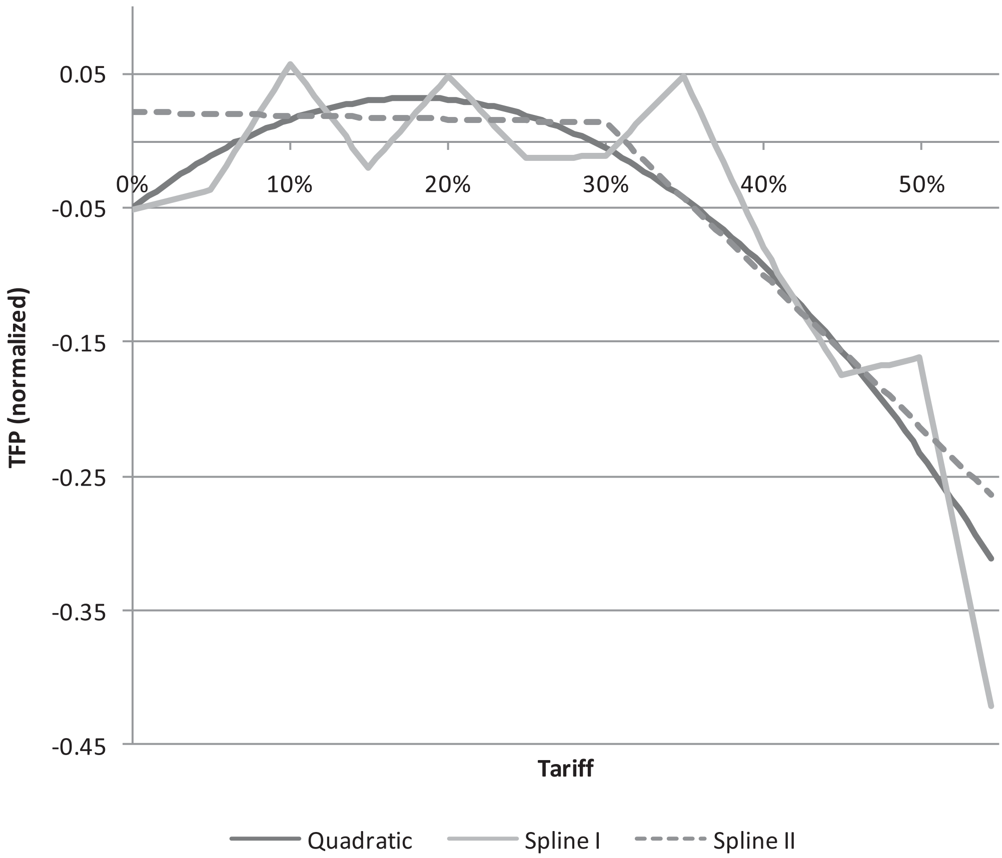


Figure 3: Predicted TFP based on Nonlinear Specifications in Table 8, col. (2), (4) & (5)



Note: Predicted TFP is normalized to zero at sample means of the explanatory variables.