Journal of International Economics 36 (1994) 53–73. North-Holland

Productivity, imperfect competition and trade reform

Theory and evidence

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Received November 1990, revised version received February 1993

Research on productivity often examines the relationship between productivity increases and structural changes in an economy, such as trade policy reform. If, however, policy reforms affect the nature of competition, then the productivity changes associated with trade reform may be mismeasured. Using a panel of manufacturing firms in the Cote d'Ivoire, this paper measures changing profit margins and productivity following a 1985 trade reform. The paper also exploits cross-section differences in protection, using data on tariffs and import penetration. Ignoring the impact of liberalization on competition leads to biased estimates of the relationship between trade reform and productivity growth.

1. Introduction

The earliest arguments for gains from trade are based on the concept of allocative efficiency. In a static framework, protection is costly because resources are not allocated in areas where a country has a comparative advantage. The recent emphasis on imperfectly competitive markets in international trade creates yet another argument for gains from trade: in a protected market dominated by only a few domestic firms, trade reform increases competition.[1]

Improving the allocation of resources or curbing firms' excess market power only generates a one-time increase in growth. Yet the 'new' endoge-

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*This paper is a revised version of an earlier draft circulated as World Bank Staff Working Paper 451 in July 1990. I would like to thank David Card, Avinash Dixit, Gene Grossman, Zvi Griliches, Jim Levinsohn, Ramon Lopez, Jaime de Melo, John Newman, John Page, J. David Richardson, Duncan Thomas and Jim Tybout for helpful comments and suggestions. Extensive comments by an anonymous referee are also very much appreciated. This paper was prepared under funding from the World Bank research project 'Industrial Competition, Productive Efficiency, and Their Relation to Trade Regimes' (RPO-674-46). This research reflects the work of the individual author and does not reflect the views and policies of the World Bank or the countries it represents.

[1]For an overview of this literature, see Helpman and Krugman (1989).

SSDI 0022-1996(93)01289-Q

nous growth theories [see Grossman and Helpman (1990) for an overview] suggest that trade policies also affect long-run growth. Moving towards free trade could permanently increase growth rates by accelerating the rate of technological change. An ever-expanding global market may raise the returns to innovaton, while more advanced technology may be embodied in imported inputs. Yet Grossman and Helpman also point out that protection could accelerate growth if it shifts resources towards manufacturing and away from research in countries with no comparative advantage in R & D. In the context of these theories, the impact of trade policies on long-run growth is ultimately an empirical question.

Despite the volume of empirical work which addresses the correlation between trade and growth, efforts to measure gains from trade at the micro level have been inconclusive. In developing countries, where small domestic markets make oligopoly behavior more likely, only a few studies link trade reform with increased competition.[2] Even more striking is the lack of conclusive evidence on the linkages between trade reform and productivity growth. There is now plant-level evidence confirming a positive relationship between trade reform and efficiency for some countries [see, for example, Tybout et al. (1991)]. Yet recent overviews on the links between trade reform and productivity growth [Bhagwati (1988), Nishimizu and Page (1990), and Tybout (1992)] suggest that the debate is still unresolved.

One possible explanation for the lack of definitive results may depend on how productivity is measured. The measurement of productivity pioneered by Solow (1957) has been used extensively to analyze technology change in both developing and developed countries. Solow derived a productivity measure, referred to as total or multi-factor productivity (TFP), which depends on two assumptions: constant returns to scale and perfect competition in product markets. Yet shifts in trade policy are likely to alter the competitive environment, particularly in developing countries where domestic markets are often dominated by several firms.

Although the potential biases from assuming perfect competition have long been recognized, this paper implements a simple approach to correct estimated TFP growth for these biases. Extending a methodology pioneered by Hall (1988) and Domowitz et al. (1988) on aggregate data, this paper uses plant-level data to explore changes in market power and productivity following a 1985 trade reform in Cote d'Ivoire. The results suggest that

[2]See, for example, de Melo and Urata (1986) who compare reported price–cost margins for two census years before and after reforms in Chile. New evidence showing that import penetration lowers price–cost margins in several developing countries will be included in Roberts and Tybout (in process). Research on developed country data includes Domowitz et al. (1986), who use aggregate data to find a negative relationship between import penetration and reported price–cost margins. See also Schmalensee (1989) for an overview of the research on developed countries.

price–cost margins fell in only a few sectors following the reform. We also explore the relationship between mark-ups and trade regime using alternative measures of 'openness', such as import penetration and tariffs. Using these alternative measures does show that increased openness to trade lowers excess profits.

When productivity estimates are modified to account for changes in price–cost margins and returns to scale, there is a stronger positive correlation between trade reform and productivity. Section 2 outlines the theoretical approach and shows how ignoring the effects of liberalization on competition may lead researchers to mismeasure the effect of trade reform on productivity. Section 3 discusses trade policy changes in the Cote d'Ivoire and briefly describes the data. We present estimation results in section 4. Section 5 incorporates the findings on market power and returns to scale to derive modified TFP estimates.

2. Correcting biases in productivity measurement

Our framework extends Hall (1988) and Domowitz et al. (1988). We begin with a production function for firm i in industry j at time t:

$$Y_{ijt} = A_{jt} f_{it} G(L_{ijt}, K_{ijt}, M_{ijt}). \tag{1}$$

Output, Y_{ijt}, is produced by firm i with inputs labor, L, capital, K, and materials, M. A_{jt} is an industry-specific index of Hicks-neutral technical progress, while f_{it} is a firm-specific parameter which allows for firm-specific differences in technology. Totally differentiating (1), and dividing through by Y, we have

$$\begin{aligned} \mathrm{d}Y/Y_{ijt} = {} & (\partial Y/\partial L)(\mathrm{d}L/Y)_{ijt} + (\partial Y/\partial K)(\mathrm{d}K/Y)_{ijt} \\ & + (\partial Y/\partial M)(\mathrm{d}M/Y)_{ijt} + \mathrm{d}A/A_{jt} + \mathrm{d}f_{it}/f_{it}. \end{aligned} \tag{2}$$

The element of imperfect competition enters (2) because firms with market power do not set the value of marginal product, $P(\partial Y/\partial L)$, equal to the factor price. If we assume Cournot behavior by firms, then we can derive the first-order conditions from each firm's profit maximization and write each of the partial derivatives $\partial Y/\partial L$, $\partial Y/\partial K$, and $\partial Y/\partial M$ as follows:

$$\partial Y/\partial L_{ijt} = (w/p)_{jt} \left[\frac{1}{1 + (S_{ij}/e_j)} \right] = (w/p)_{jt} \mu_{ij}, \tag{3a}$$

$$\partial Y/\partial K_{ijt} = (r/p)_{jt} \left[\frac{1}{1 + (S_{ij}/e_j)} \right] = (r/p)_{jt} \mu_{ij}, \tag{3b}$$

$$\partial Y/\partial M_{ijt}=(n/p)_{jt}\left[\frac{1}{1+(S_{ij}/e_j)}\right]=(n/p)_{jt}\mu_{ij}. \tag{3c}$$

Factor prices are given by w (the wage), r (the rental cost of capital), and n (the price of material inputs). If firm i is not perfectly competitive, then the value of the marginal product exceeds the factor cost by some mark-up, μ. To simplify the estimation, we will assume that the mark-up only varies across sectors. This is equivalent to assuming that market shares do not vary substantially across firms within the same sector.

Substituting (3a)–(3c) into (2) and rearranging terms, we have

$$\frac{\mathrm{d}Y}{Y_{ijt}}=\mu_j\left[\frac{wL}{PY}\frac{\mathrm{d}L}{L}+\frac{rK}{PY}\frac{\mathrm{d}K}{K}+\frac{nM}{PY}\frac{\mathrm{d}M}{M}\right]+(\mathrm{d}A/A)_{jt}+\mathrm{d}f_{it}/f_{it}. \tag{4}$$

The value of wL/PY, rk/PY, and nM/PY is simply the share of each factor (labor, capital, materials) in total output. We shall denote the share of labor and materials as α_l and α_m. Under constant returns to scale, the factor shares would sum to $1/\mu$, but we will retain a general formulation and allow the sum of the factor shares to equal β/μ, where β may be less than or greater than one.[3] Rewriting (4):

$$\mathrm{d}y_{ijt}=\mu_j[\alpha_l\,\mathrm{d}l+\alpha_m\,\mathrm{d}m]_{ijt}+(\beta-1)_j\,\mathrm{d}K/K_{ijt}+\mathrm{d}A/A_{jt}+\mathrm{d}f_{it}/f_{it}. \tag{5}$$

Lower case variables y, l and m are equal to $\ln(Y/K)$, $\ln(L/K)$, and $\ln(M/K)$. The mark-up, μ, is just the coefficient on the changes in L/K and M/K, weighted by their respective shares in output.

To see how estimates of productivity change, $\mathrm{d}A/A$, could be biased due to the presence of imperfect competition, for the moment we will assume constant returns to scale ($\beta=1$), ignore the firm-specific effect, and rewrite (5) as

$$\mathrm{d}y-\alpha_l\,\mathrm{d}l-\alpha_m\,\mathrm{d}m=\phi=(\mu-1)(\alpha_l\,\mathrm{d}l+\alpha_m\,\mathrm{d}m)+\mathrm{d}A/A. \tag{6}$$

We will refer to ϕ as the 'observed' productivity measure, and to $\mathrm{d}A/A$ as the 'true' productivity change. Under perfect competition, $\mu=1$ and $\phi=\mathrm{d}A/A$. The Solow measure of productivity, $\mathrm{d}A/A$, is unbiased.

If μ is greater than 1, however, there are two possible sources of bias. First, we may get bias in estimating the *rate* of productivity change, $\mathrm{d}A/A$. If l and

[3] Say we have a production function given by $Y=AL^aM^bK^c$, where $a+b+c$ sum to β, the scale parameter. If we take logs and differentiate, we see that

$$\frac{\mathrm{d}Y}{\mathrm{d}L}\frac{L}{Y}+\frac{\mathrm{d}Y}{\mathrm{d}M}\frac{M}{Y}+\frac{\mathrm{d}Y}{\mathrm{d}K}\frac{K}{Y}=a+b+c=\beta. \tag{a.1}$$

But from our first-order conditions, $(\mathrm{d}Y/\mathrm{d}L)(L/Y)=\mu\alpha_l$, etc. so we have $\mu\alpha_l+\mu\alpha_m+\mu\alpha_k=\beta$.

	CASE A	CASE B
	$\alpha_l dl + \alpha_m dm > 0$	$\alpha_l dl + \alpha_m dm < 0$
Pre-reform $\mu > 1$	$\phi_1 > dA_1/A_1$	$\phi_1 < dA_1/A_1$
Post-reform $\mu = 1$	$\phi_2 = dA_2/A_2$	$\phi_2 = dA_2/A_2$
Net change	$\phi_2 - \phi_1 < dA_2/A_2 - dA_1/A_1$	$\phi_2 - \phi_1 > dA_2/A_2 - dA_1/A_1$
Bias	Productivity gains under-estimated	Productivity gains over-estimated

Fig. 1. Direction of bias in productivity estimates: dA/A = true productivity; ϕ = observed productivity.

m are rising (falling), then dA/A is over (under) estimated. Second, *changes* in the trend rate of growth of productivity will be mismeasured. Fig. 1 outlines the possible biases in estimating changes in the trend rate of growth of productivity. As an example, we explore the case where price–cost margins exceed one and firms have market power prior to a trade reform. In this case, the rate of observed productivity growth, ϕ, will be greater (less) than the true measure if l and m are rising (falling). If the trade reform is accompanied by a fall in market power (possibly due to increases in the perceived elasticity of demand), price–cost margins fall to unity and measured productivity will equal the true productivity measure, dA/A. However, if we are interested in comparing productivity before and after the changes in trade policy, we are likely to incorrectly assess the true change in dA/A. As illustrated in fig. 1, the direction of the bias cannot be predicted using (6).

To see how observed productivity estimates are affected by assuming constant returns to scale, let us take the perfectly competitive case ($\mu = 1$) and rewrite (5) as (6′):

$$dy - \alpha_l dl - \alpha_m dm = \phi = (\beta - 1)\, dK/K + dA/A. \tag{6'}$$

If β exceeds 1, then the technology is characterized by increasing returns to scale. Observed productivity in this case is the sum of dA/A and productivity gains from exploiting increasing returns to scale (which are positive when dK/K is rising). Under a decreasing returns to scale technology, TFP growth is equal to the exogenous productivity term dA/A *minus* the efficiency losses that occur when a firm expands production. When observed productivity, ϕ, differs from the exogenous productivity term, dA/A, due to scale effects, one could argue that this difference is not necessarily due to 'bias'. In this case, observed productivity growth reflects both changes in Hicks-neutral technical progress and the impact of scale effects on economic efficiency.

3. Trade policy in the Cote d'Ivoire

The trade regime in the Cote d'Ivoire became increasingly restrictive in the 1970s. In 1973, a major restructuring of the tariff code increased nominal tariff rates and raised levels of effective protection by implementing an escalated tariff structure. In the second half of the 1970s and in the early 1980s, quantitative restrictions and arbitrary reference prices were introduced on a wide range of imports competing with domestic manufactures. Table 1 indicates the extent of tariff protection across industrial sectors before the reform. Textiles and food-related manufacturing received the highest nominal tariff protection, ranging from 60 to 163 percent. The pattern of effective protection, which is only available for 1980, is very similar: food processing and textiles were the most protected, followed by chemicals. Quota coverage, as measured by the number of import licenses in each sector in 1982, also follows the same pattern: textile imports were the most restricted, followed by food-related products and chemicals.

During the boom years in the second half of the 1970s, the Cote d'Ivoire benefited from the surge in world coffee and cocoa prices. The increases in revenue, most of which were captured by the government, were used to promote investment and expand public spending and infrastructure. The severe macroeconomic imbalances that followed the fall in coffee prices forced the government to adopt an austerity program in 1982. The adjustment program was followed by a major trade reform introduced in mid-1984.

The trade reform was implemented in 1985 and extended in 1986 and early 1987. The reform removed quantitative restrictions and reference prices, rationalized the tariff structure, and introduced temporary tariff surcharges. Although average tariffs did decline by 30 percent (see table 1), the goal of the tariff reform was to equalize effective protection across different sectors by lowering tariffs on final goods and raising tariffs on inputs and intermediate goods. The surcharges declined over a five-year period to allow firms previously protected by non-tariff measures to adjust.

Cote d'Ivoire's nominal exchange rate is fixed in relation to the French franc at a rate which is the same for a number of franc zone African countries. When the French franc appreciated against the U.S. dollar between 1985 and 1988, the Ivorian franc became considerably overvalued in real terms. Consequently, the reform was conducted in conjunction with an environment which lowered the competitiveness of exports on world markets.

Table 1

Sample means for selected variables.

	1979–1984				1985–1987			
Sector	Output growth	Labor output growth	TFPG[a]	Tariff	Output growth	Labor output growth	TFPG[a]	Tariff
Grain processing	−0.2	5.1	−1.7	0.50	3.2	−10.1	1.4	0.48
Food processing	6.1	8.8	2.4	0.60	−2.7	−5.7	0.9	0.59
Other food	8.5	2.9	−0.5	1.63	4.0	1.5	1.3	1.33
Textiles	4.0	1.5	0.4	0.95	0.7	−2.9	−1.9	0.72
Chemicals	2.0	−5.6	0.3	0.58	12.7	3.2	3.7	0.58
Transport	−2.2	−0.8	−1.7	0.73	7.2	−6.5	1.7	0.69
Machinery	7.2	−6.5	2.7	0.49	−3.5	−2.5	4.7	0.47
Wood products	−7.2	−0.2	−0.6	0.64	−0.4	0.0	−3.6	0.46
Paper products	−0.2	2.2	−0.8	0.57	9.6	−4.6	6.2	0.56
All sectors	2.0	−1.7	0.4	1.02[b]	4.8	−2.5	1.4	0.72

[a]TFP is defined using the Tornquist index number formula, with $\mathrm{TFP}=[\ln Y_t-\ln Y_{t-1}]-[\alpha_l\ (\ln L_t-\ln L_{t-1})+\alpha_m\ (\ln M_t-\ln M_{t-1})+(1-\alpha_l-\alpha_m)\ (\ln K_t-\ln K_{t-1})]$. Labor and material shares are defined as follows: $\alpha_l=(1/2)\ (\alpha_{lt}+\alpha_{lt-1})$, $\alpha_m=(1/2)\ (\alpha_{mt}+\alpha_{mt-1})$.

[b]Average tariffs over all sectors computed using a weighted average, using firm output as weights. A value of 0.50 indicates an ad valorem tariff of 50 percent.

Although the government simulated a partial devaluation through an export subsidy scheme for manufactured exports, the first subsidy payments were delayed until mid-1986 and payments were concentrated in several large firms. The government's inability to compensate exporting firms for the real appreciation meant that the export sector was adversely affected. Consequently, we should see a fall in price–cost margins for exporting sectors in the post-trade-reform period.

To account for changes in behavior and productivity during the trade reforms in Cote d'Ivoire, (5) is modified to allow for a change in mark-ups by firms during the post-1985 period. Changes in behavior are captured by adding an interactive slope dummy to $\mathrm{d}x$ in (5). If trade reform induced a shift in the overall level of productivity growth, then we should also include an intercept dummy:

$$\mathrm{d}y_{ijt} = B_{1j}\,\mathrm{d}x_{ijt} + B_{2j}[D\,\mathrm{d}x]_{ijt} + B_{3j}D + B_{4j}\,\mathrm{d}k_{ijt} + \mathrm{d}A/A_{jt} + \mathrm{d}f_{it}/f_{it}, \quad (7)$$

where

$B_{1j} = \mu_j$,
$B_{4j} = (\beta - 1)_j$,
$\mathrm{d}x = [\alpha_l\,\mathrm{d}l + \alpha_m\,\mathrm{d}m]$,
$\mathrm{d}k = \mathrm{d}K/K$,

D is 1 for 1985–1987 and 0 otherwise. If trade policy changes led to more competitive firm behavior, the coefficient B_2 on $[D\,\mathrm{d}x]$ should be negative, reflecting the fall in mark-ups when firms are exposed to international competition. If the reform led to overall increases in productivity growth, then B_3 should be positive. The coefficient B_4 is equal to the scale parameter β minus one.

The model presented in section 2 assumes that we correctly observe capital services, K. Yet capital services may fluctuate as capacity utilization changes over the business cycle. If we observe recorded capital stock, K^*, we can model the true utilization of capital K as equal to K^*E, where E reflects changes in utilization of capacity.

Substituting K^*E for K, the estimating equation becomes

$$(\mathrm{d}y - \mathrm{d}e)_{ijt} = B_{1j}[\mathrm{d}x - (\alpha_l + \alpha_m)\,\mathrm{d}e]_{ijt} + B_{2j}[D\{\mathrm{d}x - (\alpha_l + \alpha_m)\,\mathrm{d}e\}]_{ijt}$$
$$+ B_{3j}D + B_{4j}\,\mathrm{d}k_{ijt} + \mathrm{d}A/A_{jt} + \mathrm{d}f_{it}/f_{it}. \quad (8)$$

Since we do not have estimates of capacity utilization at the firm level, we employ a measure of total energy use as a proxy. A plant's energy use is the input component most likely to vary as capacity utilization fluctuates.

The productivity term, $\mathrm{d}A/A$, can be thought of as the average rate of productivity growth for industry j, which will be captured by a constant term B_{0j}. We will decompose $\mathrm{d}f_{it}/f_{it}$ into a plant-specific constant, g_i, plus a disturbance term, u_{it}. The final estimating equation is then given by

$$(\mathrm{d}y-\mathrm{d}e)_{ijt}=B_{0j}+B_{1j}[\mathrm{d}x-(\alpha_l+\alpha_m)\,\mathrm{d}e]_{ijt}$$

$$+B_{2j}[D\{\mathrm{d}x-(\alpha_l+\alpha_m)\,\mathrm{d}e\}]_{ijt}$$

$$+B_{3j}D+B_{4j}\,\mathrm{d}k_{ijt}+g_i+u_{it}. \tag{9}$$

Data

The firm data are taken from the Banque de Donnees Financieres (BdDF), which is instructed to gather annual information on all industrial firms. The number of firms in individual years ranges from around 250 in the 1970s to nearly 500 in the mid-1980s. Although the coverage of the industrial sector is incomplete (informal enterprises are excluded and small formal firms are under-represented), the BdDF covers almost all large and medium-sized formal manufacturing enterprises. We chose our sample of 246 firms by selecting out those enterprises with a complete time series. Although we include firms which were only present during part of the 1979–1987 period, we exclude firms which had missing values between their entry and exit dates.

We estimate (9) using our panel of 246 Ivorian firms during the period 1979–1987. The approach requires data on real output, capital stock, labor and material inputs, and the shares of labor and materials in total output. Total sales and material inputs were deflated by two-digit sectoral level price deflators to obtain a real output and materials series. We also calculated a material input price deflator based on input–output tables for each of the sectors, but the estimation results were unaffected and are not reported here. Real capital stock was constructed using the perpetual inventory method.[4]

[4]Capital stock was calculated in two steps. First, for those firms that reported across the entire sample period, we used the perpetual inventory method. Real capital stock in period t is defined in eq. (a.1): $K_{it}=(1-d)K_{i,t-1}+I_t$. As a benchmark, we used 1976 capital stock for each firm and then added real investment while accounting for depreciation. Real investment was computed by deflating nominal investment by sector-specific investment price deflators. To construct a base year real capital stock for the remaining sample of firms which entered after 1976, we first constructed a capital stock price deflator (KPD) using data on firms that were present in all years:

$$KPD_{jt}=\frac{\sum_{i=1}^{n} K_{ijt}}{\sum_{i=1}^{n} NK_{ijt}}. \tag{a.2}$$

KPD_{jt} is the capital stock deflator for sector j in year t. It was constructed using the ratio of the real capital stock computed in (a.1) to the nominal capital stock (NK) reported by firms that were present in all years. The real base year capital stock for a firm entering in year t is then given by $K_{ijt}=(KPD_{jt})\,(NK_{it})$, where t is the base year capital stock for firm i. For subsequent years, real capital stock is then computed using eq. (a.1).

The total number of employees for each firm was used as a measure of labor input.[5] Since there are only several firms in some of the more disaggregated sectors surveyed by the BdDF, we aggregated our firm sample into nine sectors: grain processing, food processing, other food, textiles, chemicals, transport, machinery, wood, and paper products. Sample means by sector for selected variables are given in table 1. Output growth during the pre-reform period averaged 24 percent, increasing to 4.8 percent during the reform. When trade reforms were introduced in 1985, the economy experienced a period of growth, but 1986 and particularly 1987 were recessionary periods. The burden of the adjustment appears to have fallen disproportionately on the labor force, with the annual average growth of employees falling from -1.7 to -2.5 percent. One shortcoming of the labor input variable is that it measures the number of permanent employees hired by the firm, but does not include information on temporary workers who may have been hired to replace the permanent labor force. However, the fall in total employment by the formal sector, documented in table 1, has been confirmed by others [see Lorch (1989)].

Table 1 also reports the uncorrected estimate for total factor productivity growth (TFPG), calculated using a Tornquist index number formula. Under trade reform, the adjusted measure shows productivity increased in most sectors but declines in others (food processing, textiles, and wood products). On average, productivity growth accelerated during the trade reform, rising from 0.4 percent to an annual average 1.4 percent growth.

4. Estimation

The final estimating equation, eq. (9), contains a firm-specific variable, g_i, which allows firms to exhibit different rates of technological change. One way to estimate this equation would be to include dummy variables for each firm. Excluding these firm dummy variables and estimating the equation by ordinary least squares (OLS) would yield biased estimates if any of the explanatory variables are correlated with the firm-specific effect. An alternative procedure is to redefine all the variables as deviations from the mean values for each firm over time, known as the within-group estimator. All the results are estimated using this approach.

4.1. Within estimates

The within estimates are first computed under the assumption of

[5]Unfortunately, the data set does not measure labor input in hours, which would allow us to control for variation in labor intensity across the business cycle.

constant returns to scale. These estimates are presented in table 2.[6] The coefficient B_1 should measure the extent of market power across sectors, while B_2 indicates the change in price–cost margins under trade reform. The mark-up of price over marginal cost is highest in the food sectors, ranging from 35 percent for the 'other food' category to 20 percent for grain processing. As indicated in table 1, these sectors received the highest levels of protection prior to trade reform. The only sector within the food category without a high mark-up is food processing, a sector which exports about 70 percent of total sales. These results suggest that heavily protected sectors which nevertheless were also highly export oriented were forced to behave competitively due to their participation in global markets.

The coefficient on B_2 is statistically significant in three out of the nine sectors. For two of those three sectors, food processing and textiles, the coefficient is negative. In the case of food processing, however, the fall in margins occurred in a sector which generally behaved competitively prior to the reforms. In this sector, the primary impetus behind falling margins was probably the large real appreciation of the exchange rate, which severely affected exporting sectors. In the case of textiles, however, anecdotal evidence suggests that margins also fell due to increased competition from the trade reform. The conversion of quotas to tariffs led to large-scale underinvoicing, which was particularly severe in the textile sector. For the other sectors, which fail to show significant effects of the reform on margins, it is possible that changes in the structure of protection were not dramatic enough to show up in the data. This is corroborated by the data on reported profit margins discussed later in the paper.

The coefficient on B_3 indicates the change in productivity growth during the trade policy reforms. The coefficient is positive for five of the nine sectors, but only statistically significant and positive for chemicals and the paper sector. Paper products experienced an unusual increase in growth during the trade reform period. Since productivity is typically procyclical, the statistically significant increase in productivity growth during the reform period may only partly be attributed to changes in the trade regime.

The within estimates are likely to be biased since inputs (including capital stock) and output are simultaneously determined by the firm. Table 2 also presents instrumental variables (IV) estimates. Appropriate instruments should be correlated with material and labor per unit of capital, but independent of any productivity or demand shocks affecting the firm. Natural instruments would include factor prices, such as wages or input prices. As instruments we use the nominal exchange rate, a price index for

[6]If f_i is random and uncorrelated with the explanatory variables, then estimating a fixed effect model using OLS will be unbiased but not as efficient as generalized least squares (GLS) estimation. Since the coefficients and standard errors for the GLS estimates were close to the within estimates, the GLS results are not reported here.

Table 2

Within and instrumental variables (IV) estimates.
Estimating equation: $dy = B_1(\alpha_l dl + \alpha_m dm) + B_2[D(\alpha_l dl + \alpha_m dm)] + B_3 D$.

	Within estimates				IV[a]			Hausman	Overid
Sector	β_1	β_2	β_3	R_2	β_1	β_2	β_3	test[b]	test[c]
Grain processing	1.201	−0.001	0.016	0.88	1.376	−0.421	−0.003	0.8	7.3
($N=83$)	(0.089)	(0.113)	(0.113)		(0.239)	(0.558)	(0.048)		
Food processing	1.031	−0.389	−0.041	0.72	1.076	−0.526	−0.059	0.1	6.2
($N=81$)	(0.084)	(0.138)	(0.039)		(0.227)	(0.407)	(0.062)		
Other food	1.353	0.013	0.012	0.92	1.415	−0.589	−0.013	1.8	4.0
($N=121$)	(0.049)	(0.090)	(0.023)		(0.162)	(0.473)	(0.038)		
Textiles	1.079	−0.253	−0.040	0.78	1.710	−0.903	−0.169	1.9	4.9
($N=183$)	(0.049)	(0.082)	(0.033)		(0.463)	(0.563)	(0.107)		
Chemicals	1.069	0.042	0.047	0.86	1.159	−0.051	0.031	1.2	4.2
($N=250$)	(0.036)	(0.069)	(0.022)		(0.091)	(0.377)	(0.027)		
Transport	1.158	−0.035	−0.016	0.92	1.096	−0.097	−0.007	1.2	0.7
($N=113$)	(0.049)	(0.069)	(0.027)		(0.135)	(0.209)	(0.032)		
Machinery	0.970	0.012	0.033	0.83	1.192	−3.439	−0.167	1.7	5.4
($N=230$)	(0.030)	(0.219)	(0.034)		(0.165)	(2.178)	(0.134)		
Wood products	1.137	0.080	−0.036	0.87	1.268	−0.172	0.052	0.2	6.1
($N=182$)	(0.048)	(0.070)	(0.030)		(0.298)	(0.569)	(0.048)		
Paper products	1.016	0.173	0.092	0.85	0.950	0.660	0.125	1.6	5.2
($N=115$)	(0.059)	(0.089)	(0.031)		(0.442)	(0.542)	(0.071)		
All sectors	1.078	0.020	0.022	0.84	1.106	−0.133	0.014	0.4	2.4
($N=1358$)	(0.016)	(0.028)	(0.010)		(0.087)	(0.247)	(0.018)		

[a]Instruments are D; the log of the nominal exchange rate; the log price index for energy; the log of debt; the log of sectoral wages; and D interacted with these four variables.

[b]The critical 5 percent value for the χ^2 (2) = 5.99. A higher value indicates a rejection of the null hypothesis that the OLS and IV estimates are the same.

[c]The over-identification test gives the χ^2 statistic for the hypothesis that the instruments are accepted as valid. The critical 5 percent of the χ^2 (7) = 14.1. A higher value indicates rejection of the test.

energy, the real wage calculated at the sector level, and the firm's reported debt. Exchange rates, wages, and energy prices should be correlated with input decisions but independent of any demand or productivity shocks affecting the firm. We also include the firm's stock of debt under the assumption that the firm's borrowings should be correlated with ability to expand inputs but are predetermined.

Following Bowden and Turkington (1984), we instrument the product D $\mathrm{d}X$ using a non-linear combination of the dummy D and $\mathrm{d}X$. In our case, this is just the set of variables, D, instruments for $\mathrm{d}X$, and the product of D and the instruments.[7] The instrumental variable estimates in table 2 were also tested for stability using various alternative sets of instruments. Our experience [which is confirmed by Abbot et al. (1989)] suggests that the standard instruments which are used in these types of regressions, such as GNP, are likely to be correlated with the error term and may lead to biased estimates of price–cost margins. Alternative specifications which employed GNP as an instrument in the first stage of the regression led to rejection of the χ^2 test used to test the model specification. One problem is that instruments which passed the test of exogeneity were sometimes poor explanatory variables for the endogenous variables. In some cases, the first-stage R^2 was less than 0.10.

We tested the validity of our instruments using a χ^2 test [see Sargan (1958) for an early exposition or Newey (1985)]. A regression of the residuals from the first stage regression on the instruments yields a χ^2 test of the validity of our instruments. The results, shown in table 2, suggest that in all cases our instruments are valid.

The instrumental variable coefficients in table 3 show a similar pattern to the within estimates. Mark-ups are highest for food-related and textile firms. However, due to the larger standard errors, we only reject the null hypothesis of perfect competition for two food sectors, textiles, and chemicals – the four sectors with the highest tariff and quota protection. Mark-ups generally fell during the trade reform period, as indicated by the negative coefficient on B_2. Again, however, none of the estimates, except for textiles, is statistically significant due to the large standard errors. A Hausman test comparing the within and IV estimates confirms the null hypothesis that the two sets of estimates are not statistically different from each other. In the remainder of this paper we present only the within estimates.

4.2. Relaxing the assumption of constant returns to scale

One potential source of misspecification arises from assuming that the

[7]If instead we had regressed $\mathrm{d}x$ on a set of instruments to obtain a predicted value for $\mathrm{d}x$, and then calculated the predicted $\mathrm{d}x$ multiplied by $D1$, this would have yielded a biased coefficient on the product $\mathrm{d}x$ $D1$.

Table 3

Within estimates with no restrictions on the scale technology estimating equation: $dy=\beta_1 dx+\beta_2[D dx]+\beta_3 D+\beta_4 dk_1$.[a]

Sector	β_1	β_2	β_3	β_4	F-test[b]
Grain processing	1.141	−0.019	0.016	0.936	1.4
	(0.105)	(0.105)	(0.030)	(0.055)	
Food processing	0.725	−0.172	−0.054	0.688	25.1
	(0.078)	(0.103)	(0.036)	(0.062)	
Other food	1.128	0.080	−0.010	0.855	10.0
	(0.078)	(0.083)	(0.023)	(0.045)	
Textiles	1.036	−0.182	−0.039	0.951	1.3
	(0.062)	(0.080)	(0.031)	(0.044)	
Chemicals	0.772	0.046	0.070	0.740	48.0
	(0.052)	(0.057)	(0.020)	(0.038)	
Transport	1.050	−0.008	0.032	0.939	2.2
	(0.067)	(0.064)	(0.025)	(0.041)	
Machinery	0.862	−0.007	0.030	0.900	34.2
	(0.037)	(0.179)	(0.033)	(0.026)	
Wood products	1.113	0.007	−0.036	0.945	1.9
	(0.052)	(0.066)	(0.029)	(0.040)	
Paper products	0.840	0.224	0.108	0.835	5.6
	(0.087)	(0.088)	(0.029)	(0.070)	
All sectors	0.947	0.040	0.024	0.881	70.8
	(0.020)	(0.026)	(0.009)	(0.014)	

[a] $dx=\alpha_l dl+\alpha_m dm$. $dk=d\log K$.
[b] Tests if coefficient on dk is equal to 1.

technology is characterized by constant returns, which permits us to omit dk from eq. (9). If the technology is not characterized by constant returns, then omitting dk leads to biased estimates. The extent of the bias can be computed as the product $R\ (\beta-1)$, where $\beta-1$ is the coefficient on dk and R is the coefficient of dk regressed on dx.[8] Assuming that R is negative, price–cost margins will be under-estimated with increasing returns to scale ($\beta>1$). If the technology is characterized by decreasing returns ($\beta<1$), then margins will be over-estimated.

Table 3 adds dk to both sides of eq. (5) and re-derives the estimating equation, eq. (9). The coefficient on dk is just β, the scale parameter. A test of constant returns to scale is then a test of whether $\beta=1$, while increasing (decreasing) returns will be indicated by a value of β greater than (less than) unity. As indicated in table 3, the coefficient on dk is generally less than 1, which implies that the technology is characterized by decreasing returns to

[8] See Schmidt (1976).

Table 4

Within estimates: Comparing different measures of openness.

	Original specification	Import penetration	Tariffs
β_1	0.947	1.091	0.941
	(0.020)	(0.033)	(0.022)
β_2	0.040	−0.248	0.039
	(0.026)	(0.053)	(0.025)
β_3	0.024	−0.186	0.016
	(0.009)	(0.093)	(0.021)
β_4	0.881	0.877	0.865
	(0.014)	(0.015)	(0.015)

scale.[9] As expected, this induced an upward bias in the coefficients reported in table 2. Table 3 shows the same pattern of price–cost margins, but the actual margins are lower. Although a number of previous studies have found increasing, not decreasing returns to scale, scale estimates are typically quite sensitive to functional form and estimation technique. In addition, it is not surprising that manufacturing firms in Cote d'Ivoire exhibit decreasing returns, since larger firms often are publicly owned or receive special treatment from the government.

4.3. *Alternative measures of trade policy*

One possible objection to our approach could be that the three years of trade reform in the Cote d'Ivoire are too brief a period to allow identification of either its impact on profit margins or productivity growth. A related objection could be raised about the depth of the reform. Since the reform sought to minimize effective rates of protection by raising some (zero) tariff rates and lowering others, the impact of the reform is difficult to measure using a time series of either average tariffs or import penetration. Although average tariffs fell by 30 percent (see table 1), import penetration stayed approximately constant. Nevertheless, combining time-series with cross-sector differences in import penetration or tariffs could be used to exploit differences in trade policy *across* sectors.[10]

We report the results of such estimates in table 4. Instead of interacting dx

[9]We also re-estimated the equation in several different ways to test whether our estimation procedure had biased the capital coefficient towards zero. The coefficient on capital remained unchanged when we estimated the equation in levels, suggesting that the within transformation of the data did not exacerbate any downward bias in the coefficient on capital stock.

[10]See also Harrison (1991), who examines the relationship between different measures of trade policy and productivity growth across countries.

with the time dummy to examine the impact of trade reform on margins, we interact dx with either import penetration or tariff rates. Data on import penetration are annual data at the sector level, while tariffs were calculated by sector for two points in time (1984 and 1987). Tariffs were measured using tariff revenues as a fraction of imports. Although this measure does not accurately reflect administrative tariff levels, particularly when there are discretionary and widespread tariff exemptions, it was the only measure available before and after the reform.

The results show a strong and statistically significant negative relationship between price–cost margins and import penetration. The coefficient on the interaction term, -0.248, suggests that if imports increased as a share of domestic consumption from 0 to 50 percent, then the excess of price over marginal cost would fall by 12 percent. As expected, price–cost margins rise as tariffs increase, but the relationship is only significant at the 15 percent level.

The coefficient on B_3 measures the change in total factor productivity growth when import penetration or tariffs rise. The coefficient is negative and statistically significant for import penetration and positive (but insignificant) for tariffs. This negative relationship between import penetration and productivity growth has been reported in other studies which use aggregate industry data to measure the relationship between productivity growth and import penetration [see, for example, Nishimizu and Page (1990)]. Yet using import penetration as a proxy for trade policy is problematic for several reasons. First, import penetration measures the outcome of changes in trade policy, not the policies themselves. It is quite possible that significant trade reforms could raise productivity without being reflected in import volumes. Second, import penetration does not allow us to distinguish between any positive effect of import competition on productivity growth in the long run and the fact that imports may be drawn to sectors with low productivity growth. Third, the observed relationship could also be explained by the pro-cyclical nature of productivity growth. In the short run, imports may lead to a contraction of output in domestic competing industries, leading to productivity losses. One way to address this problem is to use long-run averages. In section 5 we compute period averages to examine the relationship between corrected productivity growth estimates and import penetration.

4.4. Comparison with gross margins

To examine the impact of trade policy on competition, this paper makes two critical assumptions. First, we assume that trade policy is correctly measured. To test the robustness of our assumption, we have explored different approaches to measuring changes in trade policy – such as changes in policies over time vs. differences in import penetration across sectors.

Table 5

Gross margins and import penetration.

Sector	Margins pre-reform	Margins post-reform	Average margins over both periods
Grain processing	12.9	13.1	13.0
Food processing	13.1	13.9	13.4
Other food	23.4	26.6	24.5
Textiles	19.7	18.3	19.2
Chemicals	21.3	19.2	20.4
Transport	11.9	10.9	11.5
Machinery	14.0	19.5	14.7
Wood products	13.1	7.5	11.0
Paper products	15.2	21.0	17.7

Second, we assume that a fall in B_2 can be interpreted as a fall in margins. How valid is this assumption? The empirical literature in industrial organization is still wrestling with how to adequately measure departures from perfect competition.[11] In this subsection we examine one alternative to measuring margins, by examining changes in gross price–cost margins, defined as (revenue − variable costs)/revenue. As noted by Schmalensee (1989) and others, this measure is problematic since more capital-intensive industries or industries with high fixed costs will have higher margins.

Table 5 shows gross margins by sector. Average price–cost margins were estimated for the whole time period, as well as before and after the 1985 trade reform. In some cases the results are consistent with the estimates presented in table 2 and 3, while in other respects the estimates are quite different. Margins are highest in the other food category, textiles and chemicals, and lowest in wood, transport, food, and grain processing. The results are consistent with table 3 for other food (with high margins) and food processing (with the lowest margins). Gross margins did not fall significantly after the 1985 trade reform, which is consistent with our conclusions pointing to small average changes in price–cost margins overall. The sector which experienced the largest change in gross margins is paper products (where margins rose), which is consistent with the estimation results in table 3 showing that paper products are the only sector which exhibited statistically significant increases in margins. Despite these similarities, however, the correlation coefficient between estimates in tables 3 and 5 is almost zero, which suggests that the two sets of estimates may capture different aspects of firm behavior.

[11]See the chapters by Bresnahan and Schmalensee in the *Handbook of Industrial Organization* [Schmalensee and Willig (1989)].

5. Modified TFP estimates

Section 2 examined the bias in TFP measurement under imperfect competition and explored the impact on TFP measurement when the technology is characterized by increasing or decreasing returns to scale. Section 4 indicated that the manufacturing sector in Cote d'Ivoire is characterized by market power in some sectors and decreasing returns to scale. This section incorporates those findings to analyze the relationship between openness and modified estimates of productivity growth.

Although productivity may be estimated in a number of different ways, one standard approach is to use the Tornquist index number formula, which is a discrete approximation to the formula derived in eq. (6):

$$\mathrm{TFP} = [\ln Y_t - \ln Y_{t-1}] - [\alpha_l(\ln L_t - \ln L_{t-1}) + \alpha_m(\ln M_t - \ln M_{t-1}) + (1-\alpha_l-\alpha_m)(\ln K_t - \ln K_{t-1})], \tag{10}$$

where $\alpha_l = (1/2)(\alpha_{lt} + \alpha_{lt-1})$ and $\alpha_m = (1/2)(\alpha_{mt} + \alpha_{mt-1})$. If the mark-up factor, μ, and the returns to scale parameter, B, are incorporated in the definition, eq. (10) can be written as

$$\mathrm{TFP} = [\ln Y_t - \ln Y_{t-1}] - \mu[\alpha_l(\ln L_t - \ln L_{t-1}) + \alpha_m(\ln M_t - \ln M_{t-1}) + (B/\mu - \alpha_l - \alpha_m)(\ln K_t - \ln K_{t-1})]. \tag{11}$$

Estimates of μ were taken from table 4 to calculate revised TFP estimates before and after trade reform. The sample was also split into less 'open' and more 'open' using tariffs and import penetration. Estimates of productivity using both the original and revised definitions of TFPG were then recomputed for each of the six categories: time series (before and after reform); import penetration (low and high); tariffs (high and low). The results are presented in table 6. The last column of table 6 also reports (in parentheses) productivity estimates corrected for changes in capacity utilization. The estimates in parentheses are the discrete version of the estimating equation, eq. (9), derived earlier in the paper.

Under the assumption of perfect competition, we find that productivity growth increased from an annual average of 0.4 percent to 1.4 percent following the trade reform. If we relax the assumption of perfect competition and constant returns to scale, the gain in exogenous productivity is 20 percent higher. On average, the adjusted productivity growth now rises from 0.6 percent to 1.8 percent. The alternative estimates, which correct for capacity utilization, show an even greater increase in productivity growth of 2.2 percent, which is consistent with the coefficient B_3 in table 4.

Using tariffs as our measure of trade policy also yields the same result:

Table 6

Sensitivity of TFP growth estimates to assumption of perfect competition ($\mu = 1$) and constant returns to scale ($\beta = 1$).

Time series	$\mu = 1, \beta = 1$	$\mu \neq 1, \beta = 1$	$\mu \neq 1, \beta \neq 1$[a]
Pre-reform	0.4	0.5	0.6 (−1.0)
Post-reform	1.4	1.4	1.8 (1.2)
Difference	1.0	0.9	1.2 (2.2)
Tariffs			
High	0.2	0.2	0.4 (−0.4)
Low	1.2	1.2	1.6 (0.8)
Difference	1.0	1.0	1.2 (1.2)
Import penetration			
Low	0.5	0.5	0.7 (0.0)
High	1.0	1.1	1.4 (0.0)
Difference	0.5	0.6	0.7 (0.0)

[a]Numbers in parentheses indicate TFP growth adjusted for capacity utilization.

assuming perfect competition and constant returns to scale leads to underestimating the increase in productivity growth from a more open trade regime. Productivity growth is four times higher in sectors with lower tariff levels. Finally, using import penetration as a measure of trade policy yields mixed results. If we define productivity growth as in eq. (11), table 6 shows that productivity growth doubled in sectors with higher import penetration. However, if we correct for capacity utilization, the results show no relationship between import penetration and productivity growth.[12]

6. Conclusion

Research on productivity has often focused on the relationship between productivity increases and structural changes in an economy, such as trade policy reform. If, however, those structural changes affect the nature of competition or have scale effects, changes in total factor productivity growth may be mismeasured. In this paper we use a panel of firms from the Cote d'Ivoire to measure the relationship between productivity, market power, and trade reform.

[12]These results differ from reported estimates in table 4, which suggested that increased import penetration lowers productivity growth. Although both the estimates reported in table 4 and the estimates reported in parentheses in table 6 do correct for capacity utilization, import penetration is averaged over time and aggregated across sectors in table 6. Consequently, the short-run negative impact of import penetration on productivity disappears in the aggregated data.

The results suggest that firms in the most protected sectors – primarily food products oriented towards the domestic market – had the highest mark-ups of price over marginal cost. We also find weak evidence that price–cost margins fell between 1985 and 1987. In textiles, the fall in mark-ups suggests that increased competition from imports reduced the market power of existing firms. In the food processing sectors, however, the only significant declines were in exporting sectors, which were adversely affected by the real exchange rate appreciation. We also exploit the cross-section differences in import penetration and tariff rates across manufacturing sectors. Market power, as measured by price–cost margins, is significantly higher in sectors with lower import penetration and higher tariffs.

We then show that assuming perfect competition and constant returns to scale leads to under-estimating the gains in productivity growth following the trade reform. Two of the approaches to measuring the impact of trade policies show a positive association between more open trade policies and higher productivity growth. The time-series approach which compares behavior before and after 1985, shows that productivity growth tripled after the reform. Using tariffs as a trade policy measure shows that productivity growth was four times higher in the less protected sectors. If import penetration is used to capture changes in trade policy, however, the relationship between trade policy and productivity gains is more ambiguous.

Assessing the productivity effects of a trade reform, in contrast to relying on cross-section comparisons, is particularly useful if protection tends to be applied to inefficient sectors. Under those circumstances, time-series data are probably the best hope for disentangling the causal impact of a country's trade policies.

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