# Trade Protection as Income Protection in Poor Countries

**by**

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**Abstract**
This paper considers two pieces of empirical analysis of the effects of trade policy on incomes in poor countries. The first set of results relate to the relationship between trade policy and growth using a dynamic panel regression model with GMM estimates for data on 44 developing countries over 1980-1999. Trade policy is captured by measures of tariffs and import taxes, and the specification includes an interaction term between trade barriers and initial income levels to capture the non-linearity in the relationship. For low-income countries tariffs appear to be associated with *higher* growth, whereas only for middle-income and richer countries is there a negative impact of tariffs on growth. The second set of results is from a microeconometric study of the impact of trade protection on household income in Ghana. Tariff measures at the two-digit ISIC level are matched to Ghanaian household survey data for 1991/92 and 1998/99 to represent the tariff for the industry in which the household head is employed. The results suggest that higher tariffs are associated with higher incomes for households employed in the sector, at least in the short run. This positive effect of protection is disproportionately greater for less educated (low skilled labour) households, suggesting an erosion of income of unskilled labour households would result from trade liberalization. The conclusion considers some implications for political economy analysis of trade policy reform in poor countries.

**JEL Classification:** F10, F14, J31, O12, O50, O55

**Keywords:** Trade Protection; Growth; Household income; Cross-country; Ghana

## 1. Introduction

Increased globalisation and rapid trade liberalisation during the past two decades has inspired considerable debate on the impact of globalisation, in general, and trade liberalisation, in particular, on growth, incomes (and poverty). The standard argument, based on the Stolper-Samuelson theorem, is that global trade liberalisation would benefit developing countries because it would increase demand for relatively unskilled labour in which they are well endowed. While theory generally predicts that trade liberalisation would stimulate economic growth, in endogenous growth models protection of the domestic market can be growth-promoting (e.g. Grossman and Helpman, 1991). As Harrison (1996) points out, the endogenous growth theorists do not predict that free trade will unambiguously raise economic growth - increased competition could, for example, discourage innovation by lowering expected profits. The impact of trade policy on economic growth remains a matter of empirical testing. This paper discusses recent empirical evidence suggesting that protection may support incomes in poor countries, at least in the short-run. Cross-country evidence for 44 developing countries over 1980-99 finds a positive association between protection and growth for low-income countries, i.e. the effect of protection on growth only becomes negative beyond some income threshold. A possible reason is provided by evidence for Ghana, where household income tends to be higher if the household is employed in relatively protected sectors. This offers reasons why opposition to trade liberalisation is strong and widespread in poor countries, with implications for political economy analysis of globalisation.

Ackah and Morrissey (2007) investigate the impact of trade policy on economic growth in developing countries during the period 1980-1999, based on a dynamic panel regression model. They use three alternative policy measures: average unweighted scheduled tariffs, import taxes (as a percentage of imports, a measure of the average implicit tariff) and export taxes (as a percentage of exports). To allow for the differential effects on high- or low-income countries, their specification for growth includes an interaction term between trade barriers and initial income levels to capture the non-linearity in the relationship between trade barriers and growth. They address endogeneity concerns by employing the GMM estimator. The results provide evidence of a robust, positive link between trade policy and real per capita GDP growth, but the marginal impact of protection on growth is declining in income. The richer the country, the more likely it is that protection reduces growth (tariffs are negatively associated with growth), whereas for poor countries the more likely it is that trade protection will enhance growth (tariffs are positively associated with growth).

Cross-country studies only provide evidence on empirical patterns; country studies can help to explain why these patterns arise. While there have been a number of studies of the effects of trade and trade policy reform on incomes and poverty in developing countries, mostly in Latin America (see Harrison, 2005), there are very few detailed studies for SSA countries. Ackah *et al* (2007) contribute to the empirical literature with a study of Ghana. They illustrate how pseudo-panel econometric methods, appropriate in SSA where typically successive household surveys do not track the same households (i.e. they are repeated cross-sections rather than panels), can be employed to analyse the effects of trade policy changes using household data.

Ackah *et al* (2007) use repeated cross-section data (RCS) from the Ghana Living Standards Survey (GLSS) data against the background of trade reforms of the 1990s to gauge some of the effects of trade policy on households. The analyses include static and dynamic, linear and non-linear, levels and first-difference models to indicate that a higher industry tariff tends to be associated with higher income being earned by households affiliated to the industry, controlling for household-specific characteristics, geographic variables and industry fixed-effects. They find that this positive effect of protection is disproportionately greater for less educated (low skilled) labour households, suggesting an erosion of income of unskilled labour households would result from trade liberalization.

The remainder of this paper is organized as follows. In Section 2 we present the results of Ackah and Morrissey (2007), with evidence from a dynamic panel data model based on the GMM estimator of trade barriers and growth. Section 3 then presents the household-level results for Ghana from Ackah *et al* (2007). Section 4 concludes by drawing out some implications for the political economy of trade policy reform in low-income countries.

## 2. Trade, Income and Growth: Cross-country Evidence

Most of the cross-country empirical literature seems to support the view that trade liberalisation (or openness) leads to more rapid growth (Baldwin, 2003; Sachs and Warner, 1997a,b; Dollar and Kraay, 2001), although Harrison (1996) concludes that previous studies on the direction of causality between openness and growth have generated mixed results. In a critical review, Rodríguez and Rodrik (2001) contend that the cross-country growth regressions are fraught with various methodological shortcomings and the findings are less robust than claimed. The main criticisms are the unsatisfactory measures of openness commonly used, the problem of disentangling the effects of trade policies from other factors, and the appropriate way to address various econometric concerns. Rodrik (1999) argues that the benefits from openness are not *unconditional* but rather depend upon the availability of complementary policies and institutions - rule of law, good macroeconomic policies, adequate financial markets and functioning government institutions, implying a contingent or nonlinear relationship between openness and growth.

Consider the standard growth equation

$$\Delta y_{it} = \alpha\, y_{it-1} + \beta' \mathbf{x}_{it} + \eta_i + \lambda_t + \varepsilon_{it} \qquad (2.1)$$

where $y_{it}$ is per capita real GDP for country $i$ in period $t$, $\Delta y_{it}$ reflects the average growth rate of per capita GDP, $y_{it-1}$ is the initial per capita GDP, $\mathbf{x}$ is a vector of determinants of economic growth, $\eta_i$ represents the unobserved country-specific factors, $\lambda_t$ is a period-specific effect, $\varepsilon_{it}$ is the time-varying regression residual, and $\alpha$ and $\beta$ are parameters to be estimated. The subscripts $i$ and $t$ represent country and time period, respectively. Clearly, equation (1) is a dynamic model with a lagged dependent variable.

The term $\eta_i$ is a permanent but unobservable country-specific effect and captures the existence of other growth determinants that are not already controlled for by the vector $\mathbf{x}$ (i.e. omitted variables). It is time invariant and generally captures such cross sectional heterogeneity as differences in tastes or technology between countries. If the country-specific parameter were not included in (2.1), random country-specific fluctuations would be grouped into the regression residual $\varepsilon_{it}$. This would bias the common error term. In the presence of any correlation between the right-hand side variables and the country specific effect ($\eta_i$), estimation methods such as OLS will not be consistent.

Aside from omitted variable bias, cross-section growth regressions may suffer from endogeneity problems. Note that the determinants of growth in the vector $\mathbf{x}$ can be classified according to whether they are strictly exogenous, predetermined or endogenous. The vector $\mathbf{x}$ is strictly exogenous if it is uncorrelated with all past, present and future realisations of $\varepsilon_{it}$. However, this assumption is too restrictive and often very difficult to justify. For example, an unanticipated shock to the growth rate of an economy could have a contemporaneous effect on the rate of investment or the level of openness, thus compromising the strict exogeneity of these variables. Alternatively, it is reasonable to infer that a positive shock to economic growth in period $t-1$ will result in a higher level of openness or positively affect gross domestic investment in period $t$. Endogeneity is a particular problem in studies that relate growth to openness using trade outcome measures such as trade share of GDP. Even direct trade policy measures, such as average tariffs, are susceptible to potential endogeneity.

The possibility of endogeneity together with the presence of country specific effects correlated with some of the explanatory variables implies that estimation methods such as OLS will not be consistent. A first step in obtaining consistent estimates is to eliminate the country-specific heterogeneity. One approach is to employ the within-group estimator by taking deviations with respect to individual country means. However, when the model includes a lagged dependent variable the dynamic fixed-effects model produces estimates that are inconsistent if *N* (number of ‘individuals’, or cross section) is large relative to *T* (number of time periods), hence the fixed effects estimator is biased (Wooldridge, 2002). In this study the number of time periods is small ($T=5$) and thus the bias could be severe.

The growth equation (2.1) can be rewritten equivalently as

$$y_{it} = \alpha\, y_{it-1} + \beta' \mathbf{x}_{it} + \eta_i + \lambda_t + \varepsilon_{it} \qquad (2.2)$$

The Generalized Method of Moments (GMM) estimator proposed by Arellano and Bond (1991) relies on first-differencing to eliminate unobserved individual-specific effects ($\eta_i$), and then uses lagged values of endogenous or predetermined variables as instruments for subsequent first-differences. Thus, the GMM estimation procedure simultaneously addresses the problems of correlation and endogeneity. First-differencing (2.2) yields

$$y_{it} - y_{it-1} = \alpha\left(y_{it-1} - y_{it-2}\right) + \beta'\left(\mathbf{x}_{it} - \mathbf{x}_{it-1}\right) + \left(\varepsilon_{it} - \varepsilon_{it-1}\right) \qquad (2.3)$$

However, eliminating the country-specific effect introduces a correlation between the lagged dependent variable and the new error term. Also, as discussed above, the contemporaneous effects of growth shocks on the determinants of growth will result in the presence of endogeneity arising mainly due to the correlation between $\mathbf{x}$ and $\varepsilon_{it}$. To address the endogeneity problem, Arellano and Bond (1991) recommend using the lagged values of the explanatory variables in levels as instruments under the assumptions that there is no serial correlation in the error term $\varepsilon_{it}$ and the right-hand side variables. We follow DeJong and Ripoll (2004) in addressing the issue of endogeneity by imposing the identifying restriction that the determinants of growth (variables in the $\mathbf{x}$ vector) are predetermined. The assumption is that shocks to economic growth in period *t*-1 could affect, for example, physical investment, human capital investment, population growth, our trade policy measures or their interaction terms in period $t$. Given this assumption, an appropriate instrument for the difference is the lagged value.

We use the alternative systems estimator that estimates jointly the regression in differences with the regression in levels, as proposed by Arellano and Bover (1995) and Blundell and Bond (1998). The consistency of the GMM estimator depends on the validity of the assumption that the error term does not exhibit serial correlation and on the validity of the instruments. By construction, the test for the null hypothesis of no first-order serial correlation should be rejected under the identifying assumption that the error is not serially correlated; but the test for the null hypothesis of no second-order serial correlation, should not be rejected. We use two diagnostics tests proposed by Arellano and Bond (1991) and Blundell and Bond (1998), the Sargan test of over-identifying restrictions, and whether the differenced residuals are second-order serially correlated. Failure to reject the null hypotheses of both tests gives support to our model.

As we have a relatively small sample and wish to focus on the effects of the trade policy variables, we adopt a parsimonious specification including only those control variables that are most frequently found to be significant and robust. Our estimating equation in standard form is:

$$\ln Y_{it} - \ln Y_{it-1} = \delta_0 + \delta_1 \ln Y_{it-1} + \delta_2 \ln POP_{it} + \delta_3 INV_{it} + \delta_4 SEC_{it} + \delta_5 TPOLICY_{it} + \eta_i + \lambda_t + \varepsilon_{it} \quad (2.4)$$

where $Y_{it}$ is per capita GDP for country $i$ during period $t$, and $Y_{it-1}$ is the level of real per capita GDP in country $i$ at the start of period $t$. We implicitly assume that lagged income captures unobserved country-specific factors, especially those that change only slowly over time (such as institutions and geography). Population growth (*POP*) is intended to control for size effects, whereas investment (*INV*) and secondary enrolment (*SEC*) capture physical and human capital respectively. Equation (2.4) imposes a uniform and linear restriction on the parameter $\delta_5$; the *average* effect of trade policy on growth. If the growth effect of trade barriers is contingent on the level of development, Equation (2.4) may suffer from an un-modelled contingency in the relationship between trade barriers and growth (DeJong and Ripoll, 2004). Hence, we extend the basic specification to capture potential contingencies in the relationship between trade barriers and growth.

$$\ln Y_{it} - \ln Y_{it-1} = \delta_0 + \delta_1 \ln Y_{it-1} + \delta_2 \ln POP_{it} + \delta_3 INV_{it} + \delta_4 SEC_{it} + \delta_5 TPOLICY_{it} + \delta_6 \ln Y_{it-1} * TPOLICY_{it} + \eta_i + \lambda_t + \varepsilon_{it} \quad (2.5)$$

In equation (2.5), we allow the growth effect of trade policy to differ for countries at different stages of development by including an additional explanatory variable constructed as the product of initial income and our individual trade policy variables. The interaction term is meant to capture the dependence of the growth effect of trade barriers on income, where income is used here to proxy for overall level of development (as such it also captures 'initial conditions' including relatively fixed institutions). Evidence of a contingent relationship is provided by a significant coefficient on the interaction term.

*Empirical Evidence (Cross-Country, GMM)*

Annual data for 44 developing countries covering the period 1980-99 is used (limited availability of data on average tariff restricts the sample size). We construct an unbalanced panel by averaging the data over five non-overlapping four-year time periods, from 1980-83 through 1996-99. Each country thus has a potential maximum of five observations. Not all countries have data for all five time periods, but the use of unbalanced panels may attenuate the effect of self-selection in the sample. The final sample consists of 19 Sub-Saharan African countries, 11 Latin American countries, 7 from East Asia, 4 from South Asia and 3 from the Middle East and North Africa. The data comprise a heterogeneous group of countries in terms of size, level of income, degree of openness, population, resource endowments and so on. Detailed information on the data, results using alternative estimation methods and sensitivity analysis can be found in Ackah and Morrissey (2007).

The variables included in the model are widely accepted in the empirical growth literature as core determinants of growth. The log of real GDP per capita at the beginning of each 4-year period ($\ln Y_{t-1}$) is included to capture initial country-specific effects or convergence effects. If initial income captures convergence the expected sign is negative but if it captures country-specific initial conditions the sign could be positive. The coefficient on population growth (*POP*) is expected to carry a negative sign; *ceteris paribus*, more rapidly growing populations imply lower per capita income growth. The coefficients on investment share of GDP (*INV*) and human capital (*SEC*) are expected to be positive. We employ two alternative measures of trade policy (*TPOLICY*) – average (unweighted) scheduled tariffs (*TARIFF*) and import taxes as a percentage of imports (*MTAX*).

Tables 2.1 and 2.2 report coefficient estimates obtained from the growth regressions where we measure trade policy (*TPOLICY*) by average tariff (*TARIFF*) and import tax (*MTAX*) respectively (comparable, albeit weaker, results are obtained using export tax, see Ackah and Morrissey, 2007). In each of the tables, the first column is specification (2.4) where the relationship between trade barriers and growth is treated as linear – the average growth effect of trade policy. The second column is the non-linear specification (2.5) with an interaction term between policy and initial income. Results are based on the efficient system GMM estimator, for which the major diagnostic tests are reported.

The most robust variables in our regressions are investment (with the expected positive effect), the growth rate of population (expected negative effect) and human capital (expected positive effect). Ackah and Morrissey (2007) only find statistically significant effects for trade policy measures when using the GMM estimator. In line with the literature on trade restrictions and growth, it is difficult to find a consistent effect of trade barriers on growth in a linear specification of a cross-country regression, i.e. results are sensitive to data, specification and estimation technique.

In Table 2.1 (first column) *TARIFF* enters negatively but the estimated coefficient is not significantly different from zero; *MTAX* however, enters Table 2.2 negatively and is statistically significant. The results provide evidence of a globally ambiguous relationship between trade barriers and growth. When *TPOLICY* alone is introduced into the growth regression it has inconsistent signs, suggesting that it is sensitive to how trade policy is measured. It is reasonable to expect the growth effect of trade policy to differ for rich and poor countries. The results in the second column of Tables 2.1 and 2.2 explicitly allow the impact of trade policy to differ across countries in different income groups (i.e. at different stages of development), and suggest that indeed the relationship is different above and below a certain income threshold.

**Table 2.1**

***TARIFF* and Growth in Developing Countries (1980-1999): Dependent Variable is $\ln Y_{it} - \ln Y_{t-1}$**

| | LINEAR | INTERACTION |
|---|---|---|
| | [2.4] SYS-GMM | [2.5] SYS-GMM |
| $\ln Y_{t-1}$ | -0.0290*** | 0.0095 |
| | (0.0071) | (0.0074) |
| *POP* | -0.0474*** | -0.0368*** |
| | (0.0068) | (0.0112) |
| *INV* | 0.0088*** | 0.0093*** |
| | (0.0004) | (0.0005) |
| *SEC* | 0.0019*** | 0.0020*** |
| | (0.0005) | (0.0005) |
| *TARIFF* | -0.0000 | 0.0134*** |
| | (0.0003) | (0.0018) |
| *TARIFF** $\ln Y_{t-1}$ | | -0.0021*** |
| | | (0.0003) |
| Constant | 0.0903* | -0.2099** |
| | (0.0462) | (0.0786) |
| *Period Dummies:* | | |
| 1984-87 | -0.0147*** | -0.0027 |
| | (0.0047) | (0.0122) |
| 1988-91 | 0.0092*** | 0.0141* |
| | (0.0033) | (0.0073) |
| 1992-95 | -0.0324*** | -0.0193*** |
| | (0.0048) | (0.0051) |
| Sargan Test | [0.475] | [0.901] |
| $1^{st}$-order serial correlation | [0.084] | [0.090] |
| $2^{nd}$-order serial correlation | [0.538] | [0.653] |
| Observations | 136 | 136 |

*Notes:*

1. Standard errors in parentheses and *p*-values in brackets, *; **; and *** denote significant at 10%; 5%; and 1% respectively.
2. The Sargan test is for the validity of the set of instruments.
3. The tests for $1^{st}$ (m1) and $2^{nd}$ (m2) - order serial correlation are asymptotically distributed as standard normal variables (see Arellano and Bond, 1991). The *p*-values report the probability of rejecting the null hypothesis of serial correlation, where the first differencing will induce (MA1) serial correlation if the time-varying component of the error term in levels is a serially uncorrelated disturbance.

**Table 2.2**

***MTAX* - and Growth in Developing Countries (1980-1999): Dependent Variable is $\ln Y_{it} - \ln Y_{t-1}$**

| | LINEAR | INTERACTION |
|---|---|---|
| | [2.4] SYS-GMM | [2.5] SYS-GMM |
| $\ln Y_{t-1}$ | -0.0025** | 0.0222*** |
| | (0.0012) | (0.0043) |
| *POP* | -0.0437*** | -0.0299*** |
| | (0.0065) | (0.0050) |
| *INV* | 0.0097*** | 0.0096*** |
| | (0.0005) | (0.0007) |
| *SEC* | 0.0005*** | 0.0012*** |
| | (0.0002) | (0.0002) |
| *MTAX* | -0.0009** | 0.0177*** |
| | (0.0004) | (0.0023) |
| *MTAX** $\ln Y_{t-1}$ | | -0.0026*** |
| | | (0.0004) |
| Constant | -0.0529** | -0.3101*** |
| | (0.0208) | (0.0392) |
| *Period Dummies:* | | |
| 1984-87 | -0.0098 | |
| | (0.0101) | |
| 1988-91 | 0.0118* | 0.0162*** |
| | (0.0066) | (0.0060) |
| 1992-95 | -0.0263*** | -0.0151** |
| | (0.0023) | (0.0060) |
| 1996-99 | | 0.0118* |
| | | (0.0060) |
| Sargan Test | [0.722] | [0.769] |
| 1st-order serial correlation | [0.068] | [0.085] |
| 2nd-order serial correlation | [0.581] | [0.751] |
| Observations | 132 | 132 |

*Notes*: As for Table 2.1.

The relationship observed between trade policy and growth is influenced by including the interaction term: *TPOLICY* (regardless of how it is measured) enters consistently with a positive and statistically significant coefficient, but the interaction term is significantly negative in all cases. These results imply that the impact of trade barriers on growth is a function both of the level of restriction and of the level of income. Note also that the coefficient on lagged income becomes positive (albeit insignificant in Table 2.1), which we discuss below. From equation (2.5), the derivative of growth with respect to trade policy is calculated as

$$\frac{\partial GROWTH_{it}}{\partial TPOLICY_{it}} = \delta_5 + \delta_6 \left( \ln Y_{it-1} \right) \qquad (2.6)$$

This implies that the effect of a change in *TPOLICY* on *GROWTH* depends on the value of the conditioning variable, the logarithm of initial GDP per capita ($\ln Y_{it-1}$).We know from the fact that the coefficient on the interaction term is negative that the positive effect of trade barriers declines as the level of income increases. We illustrate this in Figures A1 and A2 appended, which plot the impact of a marginal change in protection on growth against real GDP per capita for our sample. For the alternative trade policy measures, the marginal effect of protection changes from positive to negative as income increases beyond the threshold level of GDP per capita. Focusing on *TARIFF*, Figure A1 reveals a threshold at the level of income equivalent to approximately $590 per capita (in constant international prices, base year 1985), above which the relationship between protection and growth is negative and below is positive.[1] Therefore, in principle, trade protection retards growth and liberalization is growth-promoting once a country has reached the threshold level of GDP per capita. A corollary is that trade liberalization will not, in general, have an unambiguous effect on growth. Trade liberalisation seems to offer the possibility of achieving faster growth only in relatively richer countries.

Sachs and Warner (1997b) offer other explanations for the sign and significance of the coefficient on the interaction term. Based on a static cross-sectional model with interaction between openness and initial income, the authors conclude that higher openness facilitates convergence; such that more open economies grow faster than closed economies. This conclusion is based on the estimated positive coefficients on both openness and the openness-initial income interaction term. In contrast, while our estimated ‘average’ coefficient on initial income is largely negative and significant (confirming the conditional convergence hypothesis), in the specifications where initial income is interacted with trade policy the estimated coefficient on initial income turns positive and significant (for *MTAX* but not *TARIFF*), implying divergence. A poor country with higher import taxes will tend to grow better than a poor country with lower import taxes, whereas a rich country with higher import taxes will tend to grow slower than a rich country with lower import taxes. The process of (non-)convergence seems to be determined, in part, by the trade regime - closed economies diverge more slowly than open economies.

[1] All the SSA countries in our sample (except South Africa, Botswana, Mauritius, Zimbabwe, Cote d’Ivoire and Congo Republic), Bangladesh, Nepal, India and Nicaragua were below this threshold level during the period 1996-99. When *MTAX* is used as our preferred measure of protection (Figure A2) the threshold level of per capita income increases to $905. When *XTAX* is used instead the threshold level of per capita income increases further to $1,167. In both cases, all the SSA countries in our sample (except South Africa, Botswana and Mauritius) fell below the relevant threshold level during the period 1996-99.

## 3 Protection and Household Income: Ghana

Theories on the impact of trade on labour income (or wages) in developing countries were traditionally based on the Stolper-Samuelson theorem, which suggests that international trade will lead to a rise in the relative returns of the abundant factor; unskilled labour in the case of developing countries. These expected gains are conditional on a number of assumptions - including free mobility of labour, given technology and perfect competition – which may not be valid for poor countries. Trade liberalization could reduce the wages of unskilled labour even in a labour abundant country if one or more assumptions are relaxed (Davis, 1996). In a world of many factors and many goods, a poor country might no longer have a comparative advantage in producing unskilled intensive goods. If a poor country has large supplies of non-labour factors of production (like land or mineral resources), trade liberalization may not benefit the labour-intensive sectors.

The specific factor and the Ricardo-Viner models have become the natural alternative to the Heckscher–Ohlin model and the associated Stolper–Samuelson theorem. According to these models workers may gain from trade reforms depending on which sectors (import-competing or exporting) they are attached to. The models focus on the short- to medium-run and assume imperfect factor mobility with one factor mobile across sectors while the other is taken to be sector-specific. With these assumptions the models predict a positive association between protection and returns to factors of production (e.g. wages). Protection reduces imports and reduced imports increase labour demand, which in turn increases wages. When the price of a good falls following trade liberalisation the model predicts that the factor specific to the sector that experienced a price reduction loses while the other factor gains in real terms. Households affiliated to the industries that experience large tariff reductions would see a decline in their incomes relative to the economy-wide average income, while households attached to other (competitive) industries would gain in comparison.

Given the apparent ambiguity in the theoretical literature discussed above the relationship between trade liberalization and incomes is ultimately an empirical matter. The non-availability (or scarcity) of panel data sets in developing countries is one of the major obstacles hampering such analysis in these countries as one cannot control for household characteristics. As we have repeated household surveys, we can apply techniques to analyse pseudo panels constructed from repeated cross sections to exploit some of the attractive features of panel data analysis such as the ability to control for household-specific effects and unobserved heterogeneity (Deaton, 1985). Two sources of data for Ghana are used to assess the impact of trade policy on household welfare during the 1990s: the GLSS conducted in 1991/92 and 1998/99 and Most Favoured Nation (MFN) tariff data for years close to the two household surveys. The latter provides tariffs, our preferred measure of trade policy, for 1993 and 2000 at the two-digit ISIC level to represent average industry-level tariffs. While this data is deficient, it should reasonably capture the distribution of tariffs and changes across sectors. This gives tariffs for 26 industries per year, of which 19 are in the traded-goods sector and seven in the non-traded sector (assumed to face zero tariffs). Our sample is restricted to households with heads aged between 18-64 inclusive, employed in any sector (tradable or non-tradable); non-working households are excluded. Each of the selected households is mapped on to one of the 26 sectors according to the sector of main employment of the household head. These exclusion restrictions leave us with a sample of 3350 and 4484 households from GLSS 3 and GLSS 4 respectively.

Among the household-level variables, we consider: demographic variables, variables relating to educational attainment, household size, linear and quadratic terms in the age of the head of the household (to capture possible life-cycle effects). We include agro-climatic zones in our model as dummy variables to control for the effects of agro-ecological zone characteristics on household welfare (this may also help capture seasonal effects due to the timing of the interview). Doing so allows us to gauge the effects of the other determinants on household income independent of the effect of agro-climatic conditions on the household. To ascertain whether there were any significant changes in household income between the two periods, we introduce a survey-year dummy, *GLSS4*. Furthermore, we allow for sectoral heterogeneity by including a dummy for households located in urban sectors, *Urban*. Using the information on the highest qualification obtained, we define five education indicators: No Education, Basic Education, Secondary Education, Post-secondary Education and Tertiary Education (University Degree). For each cross section, Table 3.1 reports summary statistics of our key variables.

Ghana embarked on a massive expansion in the provision of education during the 1990s which has resulted in the increased educational attainments during the period. The proportion of households with illiterate heads (no education) fell from 32.3 percent to 28 percent, although the share of heads with basic education has remained stable at around 57 percent. The percentage of heads with tertiary education declined marginally from 0.8 to 0.6 percent, while shares with secondary and post-secondary education rose.

Over the period we observe a decrease (from 15.9 to 11.4 percent) in the share of households employed in the public sector, consistent with the public sector retrenchment. Even though food crop farming is the largest source of employment for a great majority of households, its share declined from about 40% in 1991/92 to 37% in 1998/99. On the other hand, the share of cash crop (export) farming increased by half between the two surveys, albeit only from 5% to 7%. Non-farm self-employment saw a 14% increase in its share to remain the second largest sector of employment.

Consider the unskilled as households whose head has completed basic or no education, semi-skilled as heads who have completed secondary or post-secondary and skilled as households with university graduate heads (Ackah *et al*, 2007: Tables 3 and 4). Skilled (or semi-skilled) households are largely wage earners in either the public sector (39%) or the private formal sector (19%). Even though the unskilled dominate all socio-economic groups, almost all agriculture households (about 99% of food crop farmers and 98% of export farmers) are unskilled. While the unskilled are predominantly rural (67%) the semi-skilled (73%) and skilled (55%) are largely located in urban centres. Trends in poverty in the 1990s suggest that unskilled labour households became worse off.

**Table 3.1: Summary Statistics for Ghana Household Surveys**

| | 1991/92 | | 1998/99 | |
|---|---|---|---|---|
| **Variable** | **Mean** | **Std. Dev.** | **Mean** | **Std. Dev.** |
| Income (consumption expenditure) | 1,457,110 | 1,293,483 | 1,668,206 | 1,483,357 |
| Log Income | 13.927 | 0.710 | 14.056 | 0.729 |
| Age of head | 38.169 | 9.823 | 42.281 | 10.504 |
| Age of head squared | 1553 | 767 | 1898 | 921 |
| Female-headed household | 0.304 | 0.460 | 0.308 | 0.462 |
| *Household head has -* | | | | |
| No Education | 0.323 | 0.468 | 0.280 | 0.449 |
| Basic Education | 0.574 | 0.495 | 0.578 | 0.494 |
| Secondary Education | 0.057 | 0.231 | 0.066 | 0.248 |
| Post-secondary Education | 0.035 | 0.183 | 0.066 | 0.248 |
| Tertiary Education (University) | 0.008 | 0.091 | 0.006 | 0.074 |
| Log Value of Land | 3.510 | 5.597 | 3.419 | 6.283 |
| *Economic Activity indicators* | | | | |
| Public Sector | 0.159 | 0.366 | 0.114 | 0.318 |
| Private Formal | 0.053 | 0.224 | 0.060 | 0.237 |
| Private Informal | 0.040 | 0.197 | 0.035 | 0.185 |
| Export Farmer | 0.047 | 0.211 | 0.071 | 0.257 |
| Food Crop Farmer | 0.396 | 0.489 | 0.371 | 0.483 |
| Non-farm Self-employment | 0.304 | 0.460 | 0.347 | 0.476 |
| **Observations** | **3350** | | **4484** | |

*Source:* Authors' calculation from GLSS 1991/92 and 1998/99
*Note*: The reported figures are weighted using survey weights. Values (income and land) are in constant prices of Accra in January 1999.

Figure A3 appended shows the average tariff levels across all the 19 traded sectors in 1993 and 2000. Whereas the average unweighted scheduled tariff across *all* industries declined from 17% in 1992 to 8.5% in 1999, the structure and pattern of tariff changes was not uniform across sectors. For a sizeable number of manufacturing industries (usually, sectors with relatively skilled labour) the average tariff actually increased during the 1990s. Most manufacturing sectors continued to enjoy high levels of protection with the average tariff for industry increasing by 12 percent. The agriculture and allied industries enjoyed especially high levels of protection to begin with but these are also the sectors where tariff reductions were greatest. This suggests that Ghana protected relatively unskilled, labour-intensive sectors into the early 1990s. The rapid and substantive liberalization of trade in agriculture in the 1990s was not accompanied by similar reforms in manufacturing. What is unique about the 1990s was the sudden attempt to change the structure of protection from low-skilled agriculture and relatively low-skilled manufactures to relatively high skilled sectors.

*Empirical (Pseudo-panel) Methodology*

After matching each household with the relevant industry tariff information, we examine how the standard of living measure relates to trade protection. The approach is based on modelling the natural logarithm of per adult equivalent consumption expenditure (the income measure) of survey households, adjusted for variations in prices between localities and over time. One of the key features of the recent policy reforms in Ghana has been the significant changes in the levels of import protection. Household incomes (consumption expenditures) are likely to have been affected by the cross-sector pattern of tariffs.

Consider the determinants of household income as follows:

$$\begin{aligned} \ln w_{it} &= \alpha + \beta_1 age_{it} + \beta_2 age_{it}^{\;2} + \beta_3 hsize_{it} + \beta_4 educ_{it} + \beta_5 urban_{it} \\ &+ \beta_6 ecoz_{it} + \beta_7 land_{it} + \delta_1 tariff_{jt} + f_i + \lambda_j + \gamma_t + \varepsilon_{it} \end{aligned} \quad (3.1)$$

where the dependent variable is as previously defined, $age$ is the age of household head at the time of the survey, $age^2$ is squared age, $hsize$ is the size of the household, $educ$ is education of the household head, $urban$ is a 0/1 dummy which is 1 for households in urban localities, $ecoz$ is agro-climatic zone, $land$ is the value of land owned by the household (not area of land cultivated, in order to partly account for land quality), $tariff$ is the average tariff applied to imports of product (industry) $j$ in year $t$, $f$ is the household fixed effects, $\lambda$ is the fixed effects for the household's industry affiliation, $\gamma$ is the year fixed effect and $\varepsilon$ is the error term. The subscripts $i$ and $t$ index households and survey years respectively, year fixed effects are included to absorb economy-wide shocks (such as technological change) that may affect welfare, whilst industry dummies control for sector-specific effects.

Each of the explanatory variables is likely to explain some of the differences in household welfare. However, other unmeasured or unobservable differences among households may also matter. A pooled analysis of the data based on equation (1) will be seriously flawed, in part because such analysis cannot control for unobservables and in part because it assumes that repeated observations on each household are independent. The presence of $f$ and $\lambda$ in the model implies that we need panel data to consistently estimate the parameters in the model. To address these issues, we employ the ideas in Deaton (1985) by constructing a pseudo panel from our repeated cross-section data. Following the pseudo panel method, the first extension is to take cohort averages of all variables and estimate (3.1) based on the cohort means (see Ackah *et al* 2007, Appendix C).

$$\ln \overline{w}_{ct} = \alpha + \beta_1 \overline{age}_{ct} + \beta_2 \overline{age}_{ct}^{\,2} + \beta_3 \overline{hsize}_{ct} + \beta_4 \overline{educ}_{ct} + \beta_5 \overline{urban}_{ct} + \beta_6 \overline{ecoz}_{ct} + \beta_7 \overline{land} + \delta_1 \overline{tariff}_{ct} + \overline{f}_{ct} + \overline{\lambda}_{ct} + \overline{\gamma}_{ct} + \overline{\varepsilon}_{ct} \qquad (3.2)$$

Equation (3.2) can be estimated via random- or fixed-effects estimators. The random-effects estimator generates consistent parameter estimates if the individual effects are uncorrelated with the other explanatory variables. The fixed-effects estimator is also consistent under this assumption, but is less efficient. Under the alternative hypothesis that the individual effects are correlated with other explanatory variables, only the fixed-effects estimator is consistent. To examine whether the trade policy *changes* can be directly linked to *changes* in living standards one could also estimate a differenced model based on (3.2) as an alternative econometric specification.

The consumption (income) models (3.1) and (3.2) both assume preferences to be time separable. In effect, equation (3.2) may be misspecified (dynamically) if dynamics really matter. The best solution would obviously be to directly model the dynamics; unfortunately this is very difficult without panel data. Ackah *et al* (2007) report an alternative dynamic econometric specification, introducing the lagged dependent variable as an additional regressor, and the results are consistent with those reported here.

Following the seminal work of Deaton (1985), Ackah *et al* (2007) construct a pseudo panel and track cohorts of households through the two cross-sections. Cohorts can be defined in terms of a single characteristic or multiple characteristics. The pseudo-panel forms households into cohorts based on common multiple characteristics varying by generation (age category of head), gender of head and household's region of domicile. Given the focus on (working) households with heads between the ages of 18 to 64 and we have two cross-sections that are seven years apart then for the first cross-section (1991/92) the sample only includes households whose heads are aged 18 to 57, while the second cross-section (1998/99) only includes households with heads aged 25 to 64 so that all are in the normal working span in both surveys. Note that we add seven years to the age limits as we move to the next cross-section; this allows the households to "age" over time. We used 5-year bands in defining the generational cohorts resulting in eight birth cohorts constructed for each region in each survey year. For example, the first age cohort studied here was aged 18-22 in 1991/92 and 25-29 in 1998/99. Households whose heads are of these ages and found in the relevant cross-sections are pooled to form the pseudo cohorts. Although the actual households surveyed will differ in each survey year, they will be representative of the full cohort in the population.

As we have only two cross-sections, if the cohorts contain a large number of households, the number of cohort-groups will be small and hence the cross-sectional dimension of the panel will not be large. On the other hand, to achieve a reasonably large cross-section, some of the cohort sizes are very small, and may violate the important assumption that cohort means change only as a result of changes in household characteristics rather than in cohort composition (see Verbeek and Vella, 2005). In the psudo-panel using eight 5-year age groups, 10 regions and two gender categories, there were 160 cohorts, implying a cross-section of 320 over the two surveys (given blank cells, that actual cross-section was 310). Although average cohort-size was 52, 26 cells had fewer than 25 households (mostly in the Northern and Upper Eastern or Western regions). Thus, over ten per cent of the cells may be vulnerable to composition effects, which weakens the validity of the pseudo-panel. Some support for the approach is obtained from a second pseudo-panel using 10-year age bands to yield 148 of a potential 160 cross-section observations, with mean cell size of 104 and only eight with fewer than 25 households. The results using this smaller panel were quite similar. Detail on constructing the pseudo-panel and on estimating alternative specifications (including using the underlying household data) is in Ackah *et al*, 2007.

*Econometric Results*

The results for estimates of equations (3.1) and (3.2) are reported in Table 3.2. The first column lists the results for the case where we apply conventional OLS, based on equation (3.1), to the pooled cross-sections. Columns 2 and 3 are based on the pseudo panel equation (3.2). Column 2 reports random-effects results (tests suggest RE can be accepted as preferable to fixed-effects, which does yield broadly similar results). To examine whether the trade policy changes can be directly linked to changes in living standards we also estimate the first-difference model in column 3. This specification could also mitigate the potential for any spurious correlation between tariffs and income.

Only some control variables are significant in all three estimation methods. Incomes tend to be lower in households with older heads (*Agehead*), perhaps because they tend to have less education, and in larger households (*Hsize*). Incomes tend to be higher in *Urban* households and in the second survey (*GLSS4*). The results for the education variables are not robust: more education seems to be associated with higher income in the OLS and for tertiary education only in RE estimates, but not for differenced estimates. As mentioned above, the RE model is statistically preferred, so there is some evidence that tertiary education is associated with higher incomes, and that incomes are higher in Forest and lower in Savannah regions.

The effects of protection on income are positive and significant in all regressions in Table 3.2. Holding other factors constant, the pseudo panel econometric evidence presented here suggests that income is higher in households (or cohorts) employed in protected sectors (sheltered from competition). The coefficient on *Tariff* implies that increasing protection in a particular sector raises consumption expenditures (or incomes) in that sector. The corollary that reducing tariffs in previously protected sectors lowers incomes in those sectors is equally supported by the first-difference model.

**Table 3.2: Trade Protection and Household Income: Ghana**

| | **Cross-Section** | **Pseudo Panel** | |
|---|---|---|---|
| | *Pooled OLS* | *Random Effects* | *Differenced* |
| | **(1)** | **(2)** | **(3)** |
| Agehead | -0.022*** | -0.038*** | - |
| | (0.005) | (0.011) | |
| Agehead$^2$ | 0.001*** | 0.001*** | - |
| | (0.001) | (0.001) | |
| Hsize | -0.109*** | -0.085*** | -0.096*** |
| | (0.003) | (0.014) | (0.025) |
| Urban | 0.268*** | 0.310*** | 0.332** |
| | (0.016) | (0.077) | (0.140) |
| Basic | 0.135*** | 0.103 | 0.126 |
| | (0.016) | (0.087) | (0.193) |
| Secondary | 0.360*** | 0.434 | -0.787 |
| | (0.029) | (0.293) | (0.723) |
| Post-sec | 0.344*** | 0.414 | 0.303 |
| | (0.033) | (0.311) | (0.542) |
| Tertiary | 0.768*** | 1.880** | 1.956 |
| | (0.085) | (0.892) | (1.845) |
| Land | 0.006*** | -0.009* | -0.013 |
| | (0.001) | (0.005) | (0.015) |
| Forest | 0.017 | 0.110* | 0.026 |
| | (0.015) | (0.064) | (0.128) |
| Savannah | -0.187*** | -0.227*** | 0.169 |
| | (0.019) | (0.062) | (0.350) |
| Tariff | 0.010** | 0.056*** | 0.068** |
| | (0.005) | (0.020) | (0.029) |
| GLSS 4 | 0.127*** | 0.154*** | - |
| | (0.015) | (0.047) | |
| Constant | 14.798*** | 15.818*** | 0.185*** |
| | (0.135) | (0.897) | (0.050) |
| Industry dummies | Yes | Yes | Yes |
| Observations | 7834 | 310 | 152 |
| R-squared | 0.42 | 0.74 | 0.32 |

*Note*: Robust standard errors in parentheses, * denotes significant at 10%; ** denotes significant at 5%, *** denotes significant at 1%.

To address potential concerns about the validity of these results given their static nature and the linearity (homogeneity) restriction on the coefficient of *Tariff*, Ackah *et al* (2007) present results based on a dynamic model and including *Tariff* interacted with the *Skill* dummy (not reported here). They still find robust evidence regarding the effects of tariffs on income. Households whose heads work in industries with the largest tariff reductions (mainly the agriculture and allied sectors) would tend to experience a decline of their income relative to the economy-wide average. Households employed in more protected sectors tend to have higher incomes. This implies that some of the economic rents are shared with labour, so that liberalisation could reduce incomes and potentially increase poverty (in protected sectors).

In Table 3.3 we show the three skill types of all households in our regressions, along with their actual income as reported in the data and the predicted income from the dynamic regression, and estimate how much of the variations in within-household income is explained by trade policy. Overall, the model explains reasonably well the experience of all households irrespective of the skill type. The unexplained income (residual) is negligible, ranging between 0.3% and 5.5% in absolute terms.

**Table 3.3: Contribution of Trade Protection to Household Income**

| | **1991/92** | | | **1998/99** | | |
|---|---|---|---|---|---|---|
| | **Skill Type of Household** | | | **Skill Type of Household** | | |
| | *Unskilled* | *Semi-* | *Skilled* | *Unskilled* | *Semi-* | *Skilled* |
| Actual Income (log) | 13.875 | 14.456 | 14.324 | 13.981 | 14.586 | 14.482 |
| Predicted Income (log) | 13.870 | 14.480 | 14.378 | 13.984 | 14.571 | 14.458 |
| Residual | 0.004 | -0.024 | -0.055 | -0.003 | 0.016 | 0.025 |
| Contribution of Tariffs to Income | 0.200 | 0.184 | 0.182 | 0.176 | 0.168 | 0.168 |
| Number of Observations | 3016 | 190 | 144 | 3869 | 294 | 321 |

*Note*: Figures are simple averages over all households in each skill type except tariff which is over households in traded sectors only.
*Source*: Ackah *et al* (2007), Table 7.

The first main message from Table 3.3 is that for all the households in traded sectors the contribution of protection to income is positive. Second, the contribution of tariffs to income is slightly higher for unskilled households. Without any special safety nets or complementary policies one can expect that trade liberalisation, alone, would have disproportionate negative consequences for households in this skill type, *ceteris paribus*. Finally, the results reveal, that over the period of seven years the contribution of tariffs to income has fallen for all skill types whilst average income for each skill type has increased slightly. This suggests that in the medium to long-run there appears to be a negative relationship between trade protection and income. On implication is that our main results, in table 3.2, should be interpreted as referring to short-run effects.

## 4 Conclusions and Political Economy Implications

Section 2 examined the relationship between a trade policy measures and growth. The main result is that trade protection has, on average, a robust positive effect on economic performance for low-income countries in general. Trade liberalisation thus seems to offer the possibility of achieving faster growth only in relatively richer countries. The findings suggest that studies that have sought to explain the openness-growth relationship in terms of conventional linear models may be misleading. The conventional *average* (negative) effect of trade protection tends to mask heterogeneity in the individual responses of countries. In our sample of 44 developing countries pooled over 20 years, we find overwhelming evidence of such nonlinearity. In particular, we find that the growth effects of trade barriers may vary with the level of income from positive to negative, a possibility ignored in many previous studies. Despite being an important determinant of growth, the potential benefits from trade liberalisation are not automatic and poorer countries may actually be made worse off by it.

There are a number of reasons why protection may support growth, or alternatively why liberalisation may adversely affect growth, in the poorest countries, noting that our results should be interpreted as relating to the relatively short-run. First, import-competing sectors in these countries may be relatively underdeveloped so that even if they have the potential to be competitive and efficient, they are not so at present. This has resonance with the 'East Asian strategy' of protecting some domestic sectors at the same time as promoting export sectors. Poor countries, such as in SSA, may not be implementing such a strategy coherently and effectively, but there may be a case (and there will be a lobby) for sheltering nascent domestic industries from import competition. Second, and related, given the underdeveloped nature of the economy and the inflexibility of markets, especially limited factor mobility, the adjustment costs to trade liberalisation can be high. Third, and more generally, weak institutions and unfavourable structural characteristics (e.g. export dependence on a narrow range of primary commodities) may mean that poor countries are unable to avail of the potential benefits from liberalisation (Rodrik, 1999).

Section 3 reviews the findings of Ackah *et al* (2007) that higher tariffs are associated with higher incomes for households employed in the sector, implying that some of the economic rents are shared with labour. In Ghana, it appears that labour has been able to share in the benefits from protection, perhaps because Trade Unions are reasonably strong and thus able to negotiate higher wages. Blunch and Verner (2001) analyse the importance of unions as a determinant of earnings in the Ghanaian manufacturing sector using the Regional Program on Enterprise Development (RPED) dataset. On the basis of OLS and quantile regressions, they find evidence of a positive correlation between unionization (and firm size) and manufacturing wages in Ghana. Teal (1996) also investigates this issue using the RPED data, and provides evidence of unionization as a source of rent sharing in the Ghanaian manufacturing sector. His cross-section results reveal a strongly positive and large (larger than that found in OECD countries) effect of unionization on manufacturing wages in Ghana.

As protection has been associated with higher incomes in relatively protected sectors (i.e. higher manufacturing wages), trade liberalisation could reduce incomes and potentially increase poverty, at least in the short run, but with differing effects across skill groups. We find that the positive effect of protection is disproportionately greater for low skilled labour households, suggesting an erosion of income of unskilled labour households would result from trade liberalisation. In the short-run, all households regardless of skill type would have lost out from trade liberalisation, but the poor unskilled households (because they are sector-specific and less mobile) would lose disproportionately. The results suggest that within the same sector, a trade reform may lead to differing impacts on households with similar attributes but different skills. Moreover, education emerges as the fundamental household characteristic determining the probability that a household experiences lower income, *ceteris paribus*. From a policy standpoint, we conclude that contemplating trade liberalisation without recognizing the complementary role of human capital investment may be a sub-optimal policy for the poor, at least in the short-run. Maximizing the potential long-term benefits and minimizing the short-run costs of trade liberalisation would therefore require active interventions to address the potential adjustment costs, especially for the poor.

There are some political economy implications of the results. In Ghana, as labour appears to have been able to secure higher wages, labour and employers would have been united in lobbying for protection. This suggests that it may be difficult to mobilise support for liberalisation, suggesting an explanation for why manufacturing tariffs tended to increase despite reductions in average tariffs, especially when organised labour has a shred interest with organised employers. The cross-country evidence reinforces the difficulty of mobilising support for liberalisation: lobbies in poor countries may actually be correct in assuming that protection does raise their incomes and that of the economy. Liberalisation may yield long-run benefits, but this may be insufficient compensation for those bearing the short-run costs. Opposition to trade liberalisation will remain strong unless reforms are phased and supported by appropriate complementary policies to mitigate adjustment costs.

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Figure A1: Marginal Effect of TARIFF on Growth as a Function of GDP per Capita

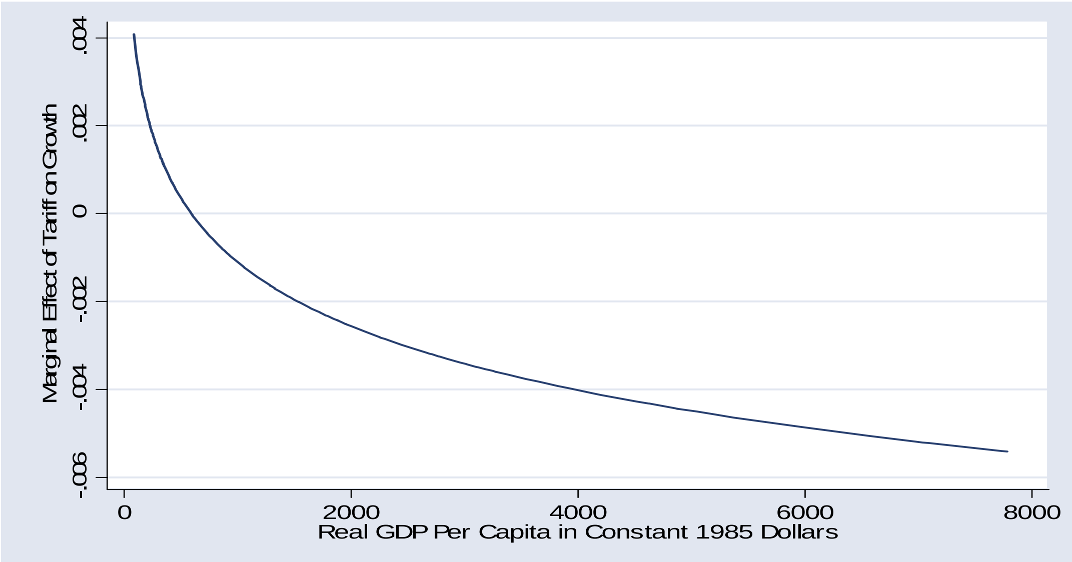
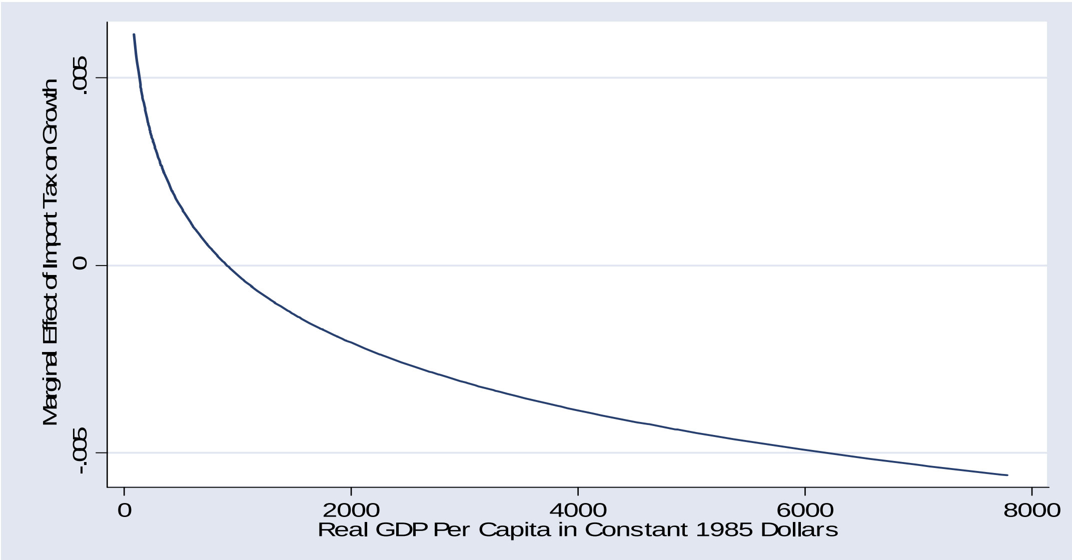
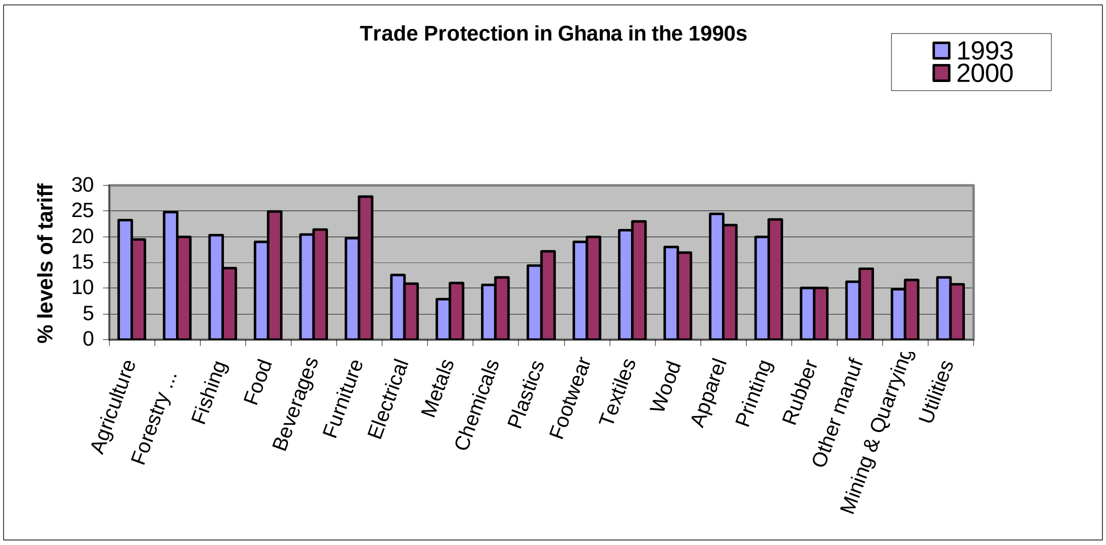


Figure A2: Marginal Effect of MTAX on Growth as a Function of GDP per Capita



## Figure A3: The Pattern of Trade Protection in Ghana during the 1990s




*Note*: These are all the 19 tradable sectors in our data. There are seven non-traded sectors with tariffs coded as zero.