Tariffs and Growth: An Empirical Exploration of Contingent Relationships

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Abstract

A large body of empirical research indicates that countries with relatively low policy-induced trade barriers tend to enjoy relatively rapid growth, ceteris paribus. In contrast, alternative theoretical models suggest that the relationship between trade barriers and growth may be contingent on the level of income. Employing a direct trade-barrier measure – ad valorem tariff rates – we find evidence of such a contingency: the marginal impact of tariffs on growth is declining in income. Moreover, evidence of a negative relationship between tariffs and growth is apparent only among the world's rich countries.
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1 Introduction

There is a widespread view, both in the economics profession and policy circles, that barriers to trade are harmful for economic growth. This view has been supported by an empirical literature in which a variety of indices of openness to international trade have been found to factor positively and significantly in cross-country growth regressions. However, in a recent review, Rodríguez and Rodrik (2001) challenge the empirical validity of the evidence this literature has produced: "... for the most part, the strong results in this literature arise either from obvious mis-specification or from the use of measures of openness that are proxies for other policy or institutional variables that

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have an independent detrimental effect on growth." [p.59][1]

A potential source of mis-specification involves the presence of an un-modeled contingency in the relationship between trade barriers and growth. The potential presence of a contingency is highlighted by a variety of theoretical models which suggest that the relationship between trade barriers and growth may be contingent on the level of development. For example, Lucas' (1988) skill-acquisition model of endogenous growth suggests that by allowing countries to establish a comparative advantage in the production of high-learning goods, the erection of trade barriers during early stages of development may enhance their long-term growth prospects. Young's (1991) learning-by-doing model carries similar implications. He shows that the growth rate of a less-developed country may decrease in going from autarky to free trade, due to the fact that comparative advantage induces these countries to specialize in goods in which the learning externality has already ceased. The models of Grossman and Helpman (1991) and Matsuyama (1992) emphasize a similar mechanism. If comparative-advantage forces prompted developing economies to concentrate resources in traditional production sectors with relatively low long-term growth prospects, openness to trade could act as a poverty trap. By temporarily shutting down such forces, trade barriers could thus enhance the growth prospects of developing economies.

Here, we do not purport to test these theoretical models directly; rather, working in the tradition of the reduced-form cross-country growth literature, we present empirical evidence of the existence of a contingent relationship between tariffs and growth. Specifically, we examine the relationship between ad-valorem tariffs and growth using a panel data set comprising 60 countries and spanning 1975 - 2000. Our baseline specification for growth features as an explanatory variable an interaction term between tariffs and initial income levels (and as an alternative, between tariffs and 1975 income rankings, as established by the World Bank).[2] This specification reveals a significant interaction effect under which the marginal impact of tariffs on growth is declining in initial income.

[1] The papers reviewed in Rodríguez and Rodrik (2000) include Dollar (1992), Ben-David (1993), Lee (1993), Sachs and Warner (1995), Harrison (1996), Edwards (1998), Wacziarg (1998), and Frankel and Romer (1999).

[2] The World Bank divides countries into four groups: low income; lower-middle income; upper-middle income; and high income. Here we represent this division using an index variable that assigns a 1 to low income countries; a 2 to lower-middle income countries; etc. Details are provided in Section 2.2.

In particular, the relationship between tariffs and growth is negative and significant among the world's rich countries, while positive (but typically insignificant, depending upon the particular estimator being employed) among the world's poor countries. This contingency pattern continues to hold under an alternative modelling strategy in which we stratify our data set into low- and high-income sub-samples (once again using World Bank rankings), and estimate separate growth-rate specifications that are linear in tariffs.

Regarding quantitative significance, under our baseline specification featuring an interaction term, we find that among countries in approximately the bottom quarter of the world's income distribution at the beginning of our sample period, a 10-percentage-point increase in tariff rates corresponds with a change in annual per capita growth of roughly 1.3 percentage points. Among countries in approximately the top quarter of the world's income distribution, the correspondence is roughly -1.6 percentage points. Stratifying the data at approximately the mid-point of the world's income distribution at the beginning of our sample period, the correspondence is estimated at roughly 0.8 percentage points among countries in the bottom half of the distribution, and at roughly -2.2 percentage points among countries in the top half of the distribution.

These estimates were obtained using the Generalized Method of Moments (GMM) dynamic panel (systems) estimator developed by Arellano and Bover (1995) and Blundell and Bond (1998). This procedure produces estimates using moment conditions arising from specifications of the underlying model expressed in both differences and levels, and is designed to account explicitly for the presence of endogenous explanatory variables and unobservable individual effects in the regression model. For comparison, we also report results obtained using three alternative estimators: cross-sectional ordinary least squares (OLS); panel seemingly unrelated regressions (SUR); and Arellano and Bond's (1991) GMM differences estimator (which employs only those moment conditions arising from specifications of the model expressed in terms of differences).[3]

As reviewed by Edwards (1993), the first generation of cross-country empirical studies on trade orientation and growth in the late 1970s focused on the relationship between export growth and

[3] Easterly and Levine (2001) employ a similar estimation strategy in establishing the influence of national policies on long-run growth. Below, we discuss pros and cons associated with the alternative estimators we employ.

GDP growth. Although these studies did not analyze trade policies directly, most of them found that "... in GDP growth regressions the importance and significance of the exports growth coefficient varies across groups of countries, casting some doubt on the desirability of pooling all these nations together in econometric analysis." [p.1383] In contrast to these earlier studies, the more recent empirical literature has not yet systematically addressed the question of whether trade policy affects growth differently across countries in different income groups.[4]

A simple illustration that the relationship between tariffs and growth is contingent on the level of income is presented in Figures 1 and 2. Figure 1 reports scatter plots and fitted relationships (estimated using partial regressions) between growth and the two tariff/initial-income interaction terms we consider.[5] In Figure 1a, the interaction is between tariffs and logged initial income; in Figure 1b, between tariffs and the World Bank's income-ranking indicator. Figure 2 plots separate relationships (again estimated using partial regressions) between growth and tariffs for the low- and high-income sample splits we consider (recall that under the sub-sample specifications, growth and tariffs are related linearly).

Figure 1 illustrates a negative relationship between growth and the interaction terms, indicating that the marginal impact of tariffs on growth is declining in initial income. In turn, Figure 2 illustrates that growth is increasing in tariffs in the low-income sub-sample (albeit insignificantly, just as is typically the case in the more formal analysis that follows), while decreasing (significantly so) in the high-income sub-sample.[6] As noted, estimates obtained using the more formal analysis

[4] In an apparent exception to this statement, Yanikkaya (2003) estimates a positive and statistically significant relationship between tariffs and growth using OLS estimation techniques, and claims this pattern to be particularly distinct among developing economies. However, as the results reported in his Table 6 (p. 82) indicate, this pattern is not robust to the removal of bias resulting from unobserved fixed effects. Moreover, the paper does not account for potential endogeneity arising from the classification of countries as developing; nor does it illustrate how tariff/growth relationships vary across income classifications. The paper does, however, provide a clear account of how alternative measures of barriers to trade relate with growth. The relationship observed between tariffs and growth is shown to be generally representative of the relationship observed for alternative measures.

[5] Partial-regression estimates are obtained in two stages. First, both the dependent variable and the isolated independent variable is projected onto the additional set of regressors under consideration (reported in Section 2.2 below). Next, the fitted dependent variable is regressed against the fitted independent variable. The figures were produced using cross-sectional regressions.

[6] These figures were produced given the exclusion of four "outlier" countries from our sample: India, Korea, Mauritius, and Papua New Guinea. The identification of these countries as outliers was made on the basis of a combination of high associated residuals and leverage statistics; details are provided in Section 3, including footnote 15. In addition, note from Figures 1 and 2 that country 43 (Nicaragua) has large associated residual values; and from

presented below confirm the presence of the contingency these figures illustrate, as well as its robustness along several dimensions.

The reminder of the paper is organized as follows. The methodology and data we employ are described in Section 2. Section 3 presents the estimation results for both the full sample of countries, and for the sample splits according to the level of development. Section 4 concludes.

2 Methodology and data

2.1 Methodology

Many different estimators have been used to evaluate the class of growth equations of interest here, and there are pros and cons associated with each (Hauk and Wacziarg, 2004, provide an excellent overview). We begin here with a brief discussion of these estimators in order to motivate the approach to estimation we adopt. Next, we discuss the method we use to explore potential contingencies in the relationship between tariffs and growth.

We analyze a panel that consists of 60 countries over the period 1975-2000; the data are split into five non-overlapping half decades. Letting the subscripts i and t represent country and time period respectively, the growth equation we estimate is given by

$$y_{it} - y_{i(t-1)} = \alpha y_{i(t-1)} + \delta' s_{i(t-1)} + \beta' c_{i(t-1)} + \xi_t + \eta_i + \epsilon_{it} \quad (1)$$

where y is the log of real per capita income observed at the beginning of the indicated time period, $s_{i(t-1)}$ is a vector of beginning-of-period state variables (stocks), $c_{i(t-1)}$ is a vector of half-decade averaged control variables (flows), ξ_t is a period-specific constant, η_i is an unobserved country-specific effect, and ϵ_{it} is an error term. Equation (1) can be rewritten as

$$y_{it} = (1+\alpha)\, y_{i(t-1)} + \delta' s_{i(t-1)} + \beta' c_{i(t-1)} + \xi_t + \eta_i + \epsilon_{it} \equiv \theta' w_{i(t-1)} + \beta' c_{i(t-1)} + \xi_t + \eta_i + \epsilon_{it} \quad (2)$$

Figure 2a, that country 33 (Kenya) also has a large associated residual. These countries were not flagged as outliers because they have low leverage values, and thus turn out to have very little influence on our results.

which makes apparent that the estimation of (1) is analogous to the estimation of a dynamic equation with a lagged-dependent variable on the right-hand side. As discussed in Caselli, Esquivel and Lefort (1996), the consistency of OLS estimators, or any variant that allows for non-spherical disturbances, depends on the assumption that the country-specific effect η_i is uncorrelated with the other right-hand-side variables. This assumption is clearly violated by (2) due to the presence of $y_{i(t-1)}$ as an explanatory variable: i.e. $E\left[y_{i(t-1)}\eta_i\right] \neq 0$. Thus, a first step in obtaining consistent estimates of this equation is to eliminate η_i.[7]

One approach to eliminating η_i that has been employed in the empirical growth literature involves its interpretation as a country-specific constant term. This interpretation motivates the implementation of a fixed-effects estimator (e.g., Islam, 1995), or the closely-related between estimator. In the former, estimates of θ' and β' are obtained using OLS applied to variables specified as deviations from period (half-decade) averages; in the latter, estimates are obtained using OLS applied directly to period averages. A second approach has involved the interpretation of η_i as a country-specific random variable that is uncorrelated with the included regressors. If the covariance structure of $\eta_i + \varepsilon_{it}$ is taken as time-invariant, θ' and β' may be estimated using feasible generalized least squares (FGLS). This is often characterized as random-effects estimation. Alternatively, if the covariance structure potentially varies over time, θ' and β' may be estimated using a seemingly unrelated regression (SUR) (see Greene, 2003, for details regarding these estimators). Under either interpretation, as Caselli et al. (1996) note, unless all explanatory variables are strictly exogenous, contemporaneous correlation between explanatory variables and error terms will remain, and inconsistency will continue to be problematic.

Alternative GMM-based approaches can be derived by taking first differences of (2), which yields

$$y_{it} - y_{i(t-1)} = \theta'\left(w_{i(t-1)} - w_{i(t-2)}\right) + \beta'\left(c_{i(t-1)} - c_{i(t-2)}\right) + \left(\xi_t - \xi_{t-1}\right) + \left(\epsilon_{it} - \epsilon_{i(t-1)}\right). \quad (3)$$

[7] A leading approach to estimating (2) is via cross-sectional OLS, wherein the panel nature of the data set is not exploited. Instead, (2) is estimated using end-of-period observations y_{iT} as dependent variables, and initial observations y_{i0} and sample averages of the additional explanatory variables as regressors. Note that the critique of Caselli et al. remains relevant in this case.

Least squares procedures cannot be used to estimate (3) because by construction, $\left(y_{i(t-1)} - y_{i(t-2)}\right)$ is correlated with $\left(\epsilon_{it} - \epsilon_{i(t-1)}\right)$. Moreover, one would like to deal with the likely endogeneity of all the explanatory variables. Arellano and Bond (1991) propose a GMM *difference* estimator that uses lagged *levels* of the explanatory variables as instruments in the estimation of (3). The estimator is based on the following identifying assumptions:

$$E\left[\epsilon_{it}\epsilon_{i(t-j)}\right] = 0 \text{ for all } j \neq 0 \tag{4}$$

$$E\left[\xi_t \epsilon_{i(t+s)}\right] = 0 \text{ for all } t\text{; and } s \geqslant 0 \tag{5}$$

$$E\left[w_{i(t-1)}\epsilon_{i(t+s)}\right] = 0 \text{ for all } t\text{; and } s \geqslant 0 \tag{6}$$

$$E\left[c_{i(t-2)}\epsilon_{i(t+s)}\right] = 0 \text{ for all } t\text{; and } s \geqslant 0 \tag{7}$$

where (4) implies that the error term ϵ_{it} is not serially correlated; (5) says that the period-specific constant ξ_t is strictly exogenous; and as indicated by (6) and (7), both the predetermined explanatory variables $w_{i(t-1)}$ and $c_{i(t-2)}$ are weakly exogenous, i.e. past stock and flow variables are uncorrelated with current and future shocks. Under these assumptions, lagged levels of the explanatory variables can be used as instruments as specified by the following moment conditions:[8]

$$E\left[\xi_{t-s}\left(\epsilon_{it} - \epsilon_{i(t-1)}\right)\right] = 0 \text{ for } s \geqslant 0,\ t = 3, \ldots.., T \tag{8}$$

$$E\left[w_{i(t-s)}\left(\epsilon_{it} - \epsilon_{i(t-1)}\right)\right] = 0 \text{ for } s \geqslant 2,\ t = 3, \ldots.., T \tag{9}$$

$$E\left[c_{i(t-s)}\left(\epsilon_{it} - \epsilon_{i(t-1)}\right)\right] = 0 \text{ for } s \geqslant 3,\ t = 3, \ldots.., T. \tag{10}$$

As discussed in Easterly and Levine (2001), this *difference* estimator has the statistical shortcoming that if the regressors in (3) are persistent, then lagged levels of w and c are weak instruments,

[8] Notice that since the control variables $c_{i(t-2)}$ correspond to half-decade averages computed between time periods $t-2$ and $t-1$, and the error term $\epsilon_{i(t-1)}$ is a function of shocks realized between periods $t-2$ and $t-1$, the levels of $c_{i(t-2)}$ are not valid instruments for the difference equation at period t. Thus, for control variables the most recent levels that are valid instruments at period t are $c_{i(t-3)}$. In contrast, the state variables $s_{i(t-2)}$ are predetermined at period $t-2$, and so are valid instruments for the difference equation at period t.

i.e. they are not highly correlated with the regressors, and so the estimated coefficients may be biased. Further, by taking differences of the original level equation (2), one loses information that speaks to the relationship between the explanatory variables and GDP growth. To overcome these issues, Arellano and Bover (1995) and Blundell and Bond (1998) developed a *systems* estimator that combines the differenced model (3) with the levels model (2). In order to be able to use lagged differences of the variables on the right-hand side of (2) as valid instruments for the regression in levels, the following identifying assumptions are introduced:

$$E\left[\eta_i\left(w_{it} - w_{i(t-1)}\right)\right] = 0 \text{ for all } t \tag{11}$$

$$E\left[\eta_i\left(c_{it} - c_{i(t-1)}\right)\right] = 0 \text{ for all } t \tag{12}$$

$$E\left[\eta_i\left(\xi_t - \xi_{t-1}\right)\right] = 0 \text{ for all } t \tag{13}$$

which imply that there is no correlation between the differences of the regressors and the country-specific effect; i.e., interactions between the country-specific effect and the regressors are stationary. Given (11) - (13), the following moment conditions can be added to those specified above in (8) - (10):

$$E\left[(\eta_i + \epsilon_{it})\left(\xi_{t-s} - \xi_{t-s-1}\right)\right] = 0 \text{ for } s = 0 \tag{14}$$

$$E\left[(\eta_i + \epsilon_{it})\left(w_{i(t-1-s)} - w_{i(t-2-s)}\right)\right] = 0 \text{ for } s = 1 \tag{15}$$

$$E\left[(\eta_i + \epsilon_{it})\left(c_{i(t-1-s)} - c_{i(t-2-s)}\right)\right] = 0 \text{ for } s = 2. \tag{16}$$

By avoiding inconsistency problems associated with OLS estimators, and weak-instrument problems associated with the difference estimator, the systems estimator seems particularly attractive in this context. But note that the move from the difference to the systems estimator also involves a cost: the adoption of the additional assumptions regarding stationarity implicit in (11) - (13), which are difficult to justify on *a priori* grounds.

An additional complication associated with the estimation of (2) involves potential measurement errors associated with regressors. Hauk and Wacziarg (2004) have studied the impact of measurement error on the performance of the estimators discussed above (with the exception of the systems estimator), explicitly in the context of the growth regressions of interest here. Their results indicate that the choice of estimators involves an additional trade-off. In the presence of measurement error, fixed-effects and difference estimators tend to underestimate $(1+\alpha)$ in (2), and thus to overestimate speed of convergence to steady state income values. In turn, these estimators tend to underestimate parameter values associated with the additional explanatory variables included in (2).In contrast, the cross-sectional OLS estimator and the panel SUR estimator both tend to provide relatively accurate estimates of speed of convergence, while overestimating the magnitude of parameters associated with the additional explanatory variables.

Lacking clear guidance regarding the choice of estimators, we follow Easterly and Levine (2001) and report results obtained from several alternative estimators: cross-section OLS, SUR, difference, and systems. SUR estimates were obtained by specifying separate regression equations for each half-decade in the sample, and allowing the covariance structure between $\eta_i + \varepsilon_{it}$ to vary across equations. In turn, regression coefficients were constrained to be constant across equations.

In addition to these estimates, between-effects, fixed-effects, and random-effects estimates were calculated but are not reported here, for two reasons. First, as characterized above, the between-effects estimator is closely related to the OLS estimator; the fixed-effects estimator is closely related to the difference estimator; and the random-effects estimator is closely related to the SUR estimator. Second, the results obtained using these additional estimators are similar qualitatively to those we report here. The additional results are available upon request.

Along with coefficient estimates obtained using the difference and systems estimators, we also report three tests of the validity of identifying assumptions they entail: Hansen's (1982) J test of over-identification; and Arellano and Bond's (1991) m_1 and m_2 tests of serial correlation properties of $\left(\epsilon_{it} - \epsilon_{i(t-1)}\right)$. Their m_1 test is of the null hypothesis of no first-order serial correlation, which should be rejected under the identifying assumption that ϵ_{it} is not serially correlated; and their m_2 test is of the null hypothesis of no second-order serial correlation, which should not be rejected.

Turning to the method we use to explore potential contingencies in the relationship between tariffs and growth, our baseline approach involves including in (2) a measure of tariffs along with an interaction term between tariffs and initial income (exact measures are described below). Evidence of a contingent relationship is provided by a significant coefficient on the interaction term. As part of an analysis of robustness, we stratify the countries in our data set into high- and low-income sub-samples (based on beginning-of-sample income levels), and estimate separate specifications of (2) that are linear in tariffs for both sub-samples. Evidence of a contingent relationship in this case is provided by a difference in tariff coefficients estimated across sub-samples. Details on the data, sample splits, etc. follow directly.

2.2 Data

As indicated above, we estimate (2) using a panel that consists of 60 countries, selected on the basis of data availability. The appendix lists all the countries in the sample, their country codes, and their classification in income groups.[9] The panel covers the period 1975-2000, and is divided into five non-overlapping half decades. The dependent variable in our sample is logged per capita real (chain weighted) GDP measured in 1980, 1985, ..., 2000.[10]

The variables used as regressors can be can be classified as stock and flow variables.[11] Stock variables are measured at the beginning of each half decade; flow variables are measured as averages over the half-decade. Stock variables consist of lagged income (so, e.g., logged GDP measured in 1975 serves as a regressor when GDP measured in 1980 is the dependent variable), and three proxies for human capital: average years of secondary school for males and females over 15 years of age from Barro and Lee's (1994) data set; and the log of life expectancy, as reported by the World Bank.

[9] We started with a sample of 74 countries for which data on real per capita GDP and tariffs were available. However, 14 of these countries were eliminated from the sample due to lack of data for other important control variables in the regressions. Also, countries with less than 4 half-decade observations were dropped from the sample due to data requirements of the Arellano and Bond estimator.

[10] Source: Penn World Tables, available at http://pwt.econ.upenn.edu. We start the time series in 1975 because the measure we use for ad-valorem tariffs, which is the main variable of interest for the paper, is available for a large number of countries only since this date.

[11] Beyond the tarrif measures we consider, we follow Barro and Lee (1994) in selecting the explanatory variables for the growth regressions. However, due to a lack of updated data, some of these regressors were not included, e.g. number of revolutions and the black market premium.

Flow variables are measured as averages over the half decade. These feature ad-valorem tariffs, measured using import duties as a percentage of imports, as reported by the World Bank.[12] They also include real private investment as a percentage of real GDP, and real government expenditures as a percentage of real GDP, both as reported in the Penn World Tables.[13]

We use three approaches to capturing potential contingencies in the relationship between tariffs and growth. Under the first, we include an additional explanatory variable constructed as the product of lagged income and tariffs. Under the second, we replace this variable with an alternative interaction term: the product of tariffs and 1975 income rankings. Following the current classification of the World Development Indicators by the World Bank, we define four income groups: 4 represents high-income countries; 3 upper-middle-income countries; 2 lower-middle-income countries; and 1 low-income countries. Since we measure income using GDP data as reported in the Penn World Tables, the country classifications we work with were obtained by mapping classification thresholds as defined by the World Bank's income measures into corresponding Penn World income measures. The resulting definitions are as follows: high-income countries are those with real per capita GDP above $11,500; upper-middle income countries those between $5,500 and $11,499; lower-middle income countries are between $2,650 and $5,499; and low-income countries those with less than $2,650. The third approach involves stratifying the data set into separate sub-samples: one that includes high- and upper-middle income countries; and one that includes lower-middle and low-income countries. Under this approach, we estimate separate specifications of (2) that are linear in tariffs for both sub-samples; evidence of a contingent relationship in this case is provided by a difference in tariff coefficients estimated across sub-samples.

Table 1 provides summary statistics for the two key variables in our analysis: tariffs and growth. Statistics are reported for the sample as a whole; for each of the four income classification groups defined as described above; and for the high- /upper-middle and low- /lower-middle income splits.

[12] Although this measure just represents average duties, it does a good job of ranking countries according their levels of protection. For a discussion on this see Rodríguez and Rodrik (2001).

[13] The measure of real government expenditures from the Penn World Tables does not exclude military expenditures. As a robustness check, we also used real government expenditures excluding military from the World Bank. Our main results do not change with either of the two measures.

All classifications are based on 1975 income rankings.

Two aspects of these statistics are of particular interest in our analysis. First is the well-known tendency for relatively rich countries to enjoy relatively rapid growth. With the important exception of the low income group, mean growth rates decrease monotonically in moving from the higher to lower income classifications: from 1.963% (standard deviation 0.633) for high-income countries to 1.459% (s.d. 2.443) for lower-middle income countries. For low-income countries, mean growth is 2.081% (s.d. 1.89). As reported in the Appendix, seven countries escaped the low-income classification in the time spanned by our sample. As part of our analysis of the robustness of our full-sample estimates, we pay close attention to the influence of the experience of these countries on our overall results. Here, we simply note that the mean growth rate of low-income countries drops to 0.45% (s.d. 2.3%) given the exclusion of these seven countries from the sample. Particularly noteworthy in this regard are Botswana and Thailand, whose growth rates averaged 4.88% and 4.77% over the sample period.

The second aspect of these statistics of particular interest here is that relatively poor countries tend to impose relatively high tariff barriers. Average tariff rates increase monotonically in moving from the high- to low-income classifications: from 2.687 (s.d. 3.635) to 17.937 (s.d. 6.895). Of course, it is inappropriate to infer a causal relationship between tariffs and economic well-being based on these statistics, but the statistics do make it tempting to view tariffs as a potential culprit behind the economic plight of poor countries. However, as the systems estimates that follow indicate, the situation is more complicated than these statistics suggest.

3 Estimation results

Estimation results are presented in Tables 2 – 9.[14] Table 2 reports coefficient estimates obtained using the full set of countries in our sample; Table 3 reports estimates obtained given the elimination of four potential outlier countries: India, Korea, Mauritius, and Papua New Guinea. These countries were singled out using a strategy advocated by Belsley, Kuh and Welsch (1980), which

[14] All estimates were computed using Stata, Version 8.0.

involves the use of the so-called DFITS statistic to flag countries associated with high combinations of residual and leverage statistics.[15] Table 4 is a companion to Table 2, wherein we report a measure of quantitative significance implied by the parameter estimates reported in Table 2; the measure reports the associated impact on growth of a ten-percentage-point increase in tariff rates. Table 5 is an analogous companion to Table 3. Table 6 reports coefficient estimates obtained using the alternative to our baseline specification under which linear growth/tariff relationships are estimated using low- and high-income sub-samples of the data. Tables 7 and 8 report on the sensitivity of our baseline estimates to the presence in the data of several alternative subsets of countries, singled out for certain unusual aspects of their growth-rate experiences (Table 7) and tariff rates (Table 8). Finally, Table 9 presents estimates designed to determine whether the relationship between tariffs and growth is contingent on variables other than income.

Coefficient estimates associated with lagged income and logged life expectancy are interpretable as elasticity measures; all other coefficients represent semi-elasticity measures. In addition, the measures of quantitative significance reported in Tables 4 and 5 were computed as follows. Let β_1 denote the coefficient associated with tariff rates in (2), β_2 the coefficient associated with the tariff/income interaction term, and INC the measure of income used to construct the tariff/income interaction term (recall that we use two alternative measures of income for this purpose: logged initial income, and the World Bank income ranking index). Then, measured on an annual percentage basis, the associated impact on growth of a 10-percentage-point increase in tariffs is given by

$$100\left(\beta_1 + \beta_2 INC\right)\left(\frac{10}{T}\right), \tag{17}$$

where T is the time gap between the observations of $y_{i,t}$ and $y_{i,(t-1)}$ in (2) (the time gap is 25 for our

[15] Letting X denote the matrix of explanatory variables under consideration, with ith row x_i containing observations on country i, and H denote the associated "hat" matrix $X(X'X)^{-1}X'$, the leverage statistic h_i for country i is the ith diagonal element of H. This is a measure of the distance of x_i from the center of mass from the other rows of X. Letting r_i denote the residuals associated with country i, the $DFITS_i$ statistic is given by $DFITS_i = r_i\sqrt{\frac{h_i}{1-h_i}}$. Following Belsley, Kuh and Welsch (1980), we identified as potential outliers countries with associated $DFITS$ statistics greater than $2\sqrt{k/n}$, where k denotes the number of explanatory variables and n the number of countries under consideration. This identification proved insensitive to the use of alternative diagnostic measures.

cross-sectional OLS estimates, and 5 for our panel estimates), and multiplication by 100 facilitates conversion to a percentage-point measurement. Standard errors associated with this measure are obtained using the delta method. Finally, we also report values of λ (annualized rates of conditional convergence, or divergence if negative) implied by the coefficient estimates associated with lagged income α; i.e., λ solves

$$1 + \alpha = \exp(-\lambda t),$$

where t is the time distance between current and lagged income.[16]

We begin with a note regarding tests of the identifying assumptions upon which our GMM estimates are based. Recall that the m_1 tests we report are of the null hypothesis of no first-order residual serial correlation; this hypothesis should be rejected (in favor of negative serial correlation) under our assumptions. With one exception (for results obtained using the GMM difference estimator in Table 6b, which should thus be interpreted with caution), this hypothesis is indeed rejected at the 5% significance level, at least. The m_2 tests are of the null hypothesis of no second-order residual serial correlation; this hypothesis should not be rejected under our assumptions. This is the case without exception, at no less than the 35% significance level. Finally, the tables report "H": Hansen's (1982) test of the over-identifying restrictions implied by our choice of instruments. The null hypothesis should not be rejected under this test, and indeed in all cases, it is not (at the 50% significance level, at least).

3.1 Convergence

We begin with a discussion of convergence estimates. While not reported in the tables, unconditionally, the full set of countries in our sample exhibit a strong tendency towards divergence (at an annualized rate of 1.37% according to GMM systems estimates). This tendency is similar to that observed among broader samples of countries (e.g., Pritchett, 1997). Regarding the conditional convergence estimates reported in the tables, these vary systematically across the estimators used to obtain them. Excepting estimates obtained using the GMM difference estimator, the variation

[16] For details regarding the link between α and λ see Barro and Sala-i-Martin (2003), section 2.6.6, p. 111.

is between 0.99% and 2% for the full-sample estimates (with and without outliers, as reported in Tables 2 and 3); 1.87% and 4.25% among low and lower-middle income countries (Table 6a); and 2.48% and 5.52% among upper-middle and high income countries (Table 6b). GMM systems estimates of the speed of convergence tend to be statistically insignificant, while the additional estimates are typically significant. To put these estimates in context, we note that, using cross-sectional OLS, Barro and Sala-i-Martin (2003, Table 12.3, p. 522) obtain a convergence estimate of 2.48% using a broader set of explanatory variables than we consider. Our full-sample OLS estimates (with and without outliers) lie within one standard deviation of this estimate.

In contrast, conditional convergence estimates obtained using the GMM difference estimator are extremely high: in the neighborhood of 9% in the full sample (Table 2), and 11% given the removal of outliers (Table 3). This is a result akin to that obtained by Caselli et al. (1996), who also obtained relatively rapid conditional convergence estimates (ranging from 5.4% to 7.9% under their specifications) using the GMM differences estimator. Caselli et al. attributed the relatively rapid convergence estimates they obtained to the removal of downward bias in OLS estimates resulting from likely positive correlation between untreated individual effects and initial income levels. However, since the GMM systems estimator also removes this downward bias, and yet delivers relatively low estimates of conditional convergence, this explanation may be tenuous. It may instead be the case that Caselli et al.'s estimates reflect a weak-instrument bias that is known to plague the difference estimator, particularly given the use of persistent explanatory variables (e.g., Blundell and Bond, 1997). It may also be the case that problems associated with measurement error are operative, as suggested by the findings of Hauk and Wacziarg (2004): recall that in the presence of measurement error, the difference estimator tends to overestimate speed of convergence, and underestimate parameters associated with additional explanatory variables. A full resolution of this issue is beyond the scope of the present analysis, but it is interesting to note that the relatively efficient GMM-systems estimator we employ here returns us closer to the range of convergence estimates traditionally associated with OLS estimators.

3.2 Tariff/growth contingencies

We now turn to relationships observed between tariffs and growth. While not reported in the tables, we begin by discussing the global relationship observed between tariffs and growth given the exclusion of a tariff/income interaction term from our baseline specifications (i.e., given a specification of (2) that is linear in the regressors, estimated without stratifying countries into high- and low-income subsets). Using the full data set, we estimate a moderate positive relationship, putting the estimated impact on growth (again measured as an annual percentage) of a ten-percentage-point increase in tariffs in the neighborhood of 0.7%; dropping the four outlier countries from the sample reduces this figure to 0.4% (the former estimate is marginally significant, the latter is insignificant). We view this finding as indicating the empirical importance of the Rodriguez-Rodrik (2000) critique: controlling for individual effects and eschewing the use of indices of trade openness, we find no evidence globally of a negative relationship between growth and the trade barrier we measure directly. As the contingencies to which we now turn indicate, a negative relationship is evident only among the relatively rich countries of the world.

Regarding contingencies, there are sixteen sets of estimates reported in Tables 2 and 3: all sixteen include a negative coefficient on the tariff/income interaction term. In only one of four cases is the estimate obtained using the GMM difference estimator significant at the 10% level, perhaps reflecting once again the bias (in this case, towards zero) that afflicts this estimator in the presence of measurement error (Hauk and Wacziarg, 2004). In contrast, the cross-sectional OLS, SUR and GMM systems estimators each yield significant coefficients at the 10% level in three out of four cases.

The associated quantitative significance of these estimates is reported for the full sample in Table 4, and given the removal of outliers in Table 5. As noted, we measure quantitative significance as the impact on growth of a ten-percentage-point increase in tariffs, calculated as in (17). Note from the tables that four sets of measures are reported for each set of parameter estimates; this reflects the use of four values of INC in (17) for each set of parameter estimates. For the specifications under which tariffs were interacted with the World Bank's income-rank index, the four values assigned to INC are simply the four index values $1 - 4$. For the specifications under which tariffs were interacted with logged initial income, INC is assigned average values of logged initial income calculated within each income group. In either case, the tables report average impacts on growth estimated within the four sets of country classifications we consider. Reported p-values in this case are associated with the null hypothesis that the estimated impacts are zero.

For the poorest countries in the world (those with index values of 1), all sixteen sets of quantitative-significance measures reported in Tables 4 and 5 are positive, with estimates ranging from 0.117% (as reported for the SUR estimator in Table 5a) to 1.38% (as reported for the GMM systems estimator in Table 5a). And for the richest countries in the world, all sixteen measures are negative, and range from -0.04% (as reported for the OLS estimator in Table 4a) to -3.476% (as reported for the GMM difference estimator in Table 5b). For the lower- and upper-middle income countries, a mixed picture emerges: the measures oscillate between positive and negative values across estimators and data samples, and rarely differ significantly from zero.

There is a notable difference across the measures reported in Tables 4 and 5 that signifies the influence of the outlier countries in the data set. In Table 4, measures obtained using the OLS and GMM systems estimators for the low-income countries differ significantly from zero, while the measures reported Table 5 are typically insignificant. In contrast, measures estimated for the high-income countries typically differ significantly from zero in Table 5, but not in Table 4. So while in both versions of the data there tends to be a distinct difference in tariff/growth relationships across income groups both statistically and quantitatively, the removal of the outlier countries results in a de-emphasis of the positive tariff/growth relationship observed among poor countries, and an emphasis of the negative relationship observed among rich countries.

A similar picture emerges when splitting the data set into sub-samples, one including low- and lower-middle income countries, the other including upper-middle and high-income countries, and estimating separate linear specifications for each. Results of this exercise are reported in Table 6.[17]

[17] The results in Table 6 were obtained given the exclusion of the four outlier countries identified above from the former sample, and in addition, Iceland and Venezuela from the latter sample. In neither case did the exclusions affect the pattern of quantitative significance evident in Table 6, and in all cases, estimated tariff coefficients were all within one standard deviation; thus we report only one set of estimates for each sub-sample here.

For the low and lower-middle income sub-sample, the tariff coefficients are all estimated as positive and insignificant. For the upper-middle and high-income sub-sample, the tariff coefficients are all estimated as negative and significant at the 10% level. Regarding quantitative significance, using the estimates produced by the systems estimator, the impact on growth of a 10-percentage-point increase in tariffs is estimated as 0.8 percentage points among the low- and lower-middle-income countries, and −2.2 percentage points among upper-middle and high-income countries.[18] Thus, the results obtained when modelling the tariff/growth contingency are robust to the use of this alternative sample-splitting methodology.

Before describing some additional checks for robustness we conducted, we pause briefly to discuss coefficient estimates obtained for the additional explanatory variables included in (2), which turn out to be quite standard. Estimated coefficients on the I/Y (G/Y) ratio are consistently positive (negative), statistically significant, and typically indicate strong quantitative effects. Life expectancy also exhibits a strong positive relationship with growth, but note from Table 6 that it tends to factor more substantially among low- and lower-middle income countries than among upper-middle and high-income countries. Another interesting difference across sub-samples concerns the importance of female schooling, which tends to factor positively and significantly among upper-middle and high-income countries, but insignificantly among low- and lower-middle income countries. Finally, note that rates of "club convergence" are approximately twice as rapid among upper-middle and high-income countries than among low- and lower-middle income countries.

3.3 Robustness checks

Beyond the robustness checks described above, we paid special attention to the potential influence on our results of several subsets of countries. The first collection of subsets features countries singled out on the basis of certain unusual aspects of their growth rate experiences during the time period spanned by our sample. Results of this exercise are reported in Table 7 for four subsets of countries.

[18] These measures are obtained by multiplying the coefficient estimate by the percentage-point change of 10, dividing by the time span between income observations (5 years), then multiplying by 100 to convert to a percentage-point measurement.

For each subset, Table 7 reports the list of countries, their 1975 and 1995 income rankings, their average tariff and growth rates measured over the entire sample period, and coefficient estimates obtained for the tariff/income interaction term given their removal from the sample. We report on only those estimates obtained using the systems estimator, but the general flavor of the exercise is consistent across estimators. For ease of comparison, Table 7 also includes the full-sample estimates reported in Table 2, and estimates obtained given the exclusion of the four outlier countries, as reported in Table 3. Beyond the four outlier countries, the additional subsets of countries singled out in Table 7 include the seven escapees from the low-income group (as originally classified in 1975) mentioned above; seven escapees from either the low- and lower-middle income groups (i.e., escapees from the sub-sample highlighted in Table 6a); and the Asian Tigers. Average growth rates among the countries within each group tend to be quite high, accounting for their mobility between rankings.

Strikingly, coefficient estimates change very little given the removal of any one of the subsets considered in Table 7. For both interaction terms, estimates obtained given the removal of each sub-sample lie within one standard deviation of the full-sample estimate. Regarding the coefficient on the tariff/initial income term, parameter estimates range from the full-sample estimate of -0.006 (s.e. 0.003) to -0.008 (s.e. 0.003), which is the coefficient obtained given the removal of the outliers highlighted in Table 3. For the tariff/income-rank term, estimates range from the reported estimate of -0.003 (s.e. 0.002) in Table 3 to the full-sample estimate of -0.005 (s.e. 0.002). An apparent explanation for the lack of leverage exerted by these sub-samples is that, while the countries they contain are typified as having low initial incomes and high subsequent growth rates, they are not uniformly associated with unusually high tariff rates. Indeed, only three have tariff rates that exceed the sample average of 9.61 by the sample standard deviation of 7.85: India (33.85), Egypt (25.61), and Botswana (19.19). And in the case of the Asian Tigers, rather than serving as a potential driving force behind our results, they actually appear as exceptions to the rule: excepting Singapore, they rank initially in the bottom half of the World income distribution, maintained low average tariff rates, and enjoyed above-average growth rates.

The second collection of subsets we examined include countries singled out due to the maintenance of exceptional tariff rates. We considered three subsets: the four high-income countries with the lowest average tariff rates; the four low-income countries with the highest average tariff rates; and the union of these two subsets.[19] The impact of removing these subsets of countries is reported in Table 8. Once again, point estimates change very little; what does change somewhat in this case is statistical significance. In particular, neither of the interaction coefficients is statistically significant given the removal of all eight countries (p-values are 0.145 and 0.188 in this case); and the coefficient on the tariff/initial income term is insignificant given the removal of the four low-income countries (its p-value is 0.152). Excepting Cameroon, the combinations of high tariffs and growth rates observed among the low-income countries highlighted here suggest their potential for leveraging our results, and indeed, India stands out as a true outlier in the sample (thus its treatment as such through the course of our analysis). However, the general pattern of results we have reported remains apparent given their exclusion from the sample.

Collectively, the results of Tables 7 and 8 suggest that the contingency we have document in the relationship between tariffs and growth does not seem attributable to the influence of a small number of exceptional countries. Rather, it seems fairly described as being pervasive through the data set as a whole.

We conclude by briefly describing three additional sensitivity checks that further underscore this point. First, we augment the regressions reported above to control for additional measures of institutional strength. We consider three indicators: the economic freedom index of The Fraser Institute, and the indices of democracy and constraints to the executive from Polity IV.[20] When

[19] There are 17 high-income countries in the sample, and 16 low-income countries; so the two subsets represent approximate quartiles of their respective distributions.

[20] The economic freedom index is taken from Gwartney et al. (2001), and is available electronically at www.freetheworld.com. This index ranks countries from 0 to 10, with 10 meaning more economic freedom. The index is a weighted average of sub-indexes in the following seven areas: size of government, structure of the economy and use of markets, monetary policy and price stability, freedom to use alternative currencies, legal structure and propery rights, international exchange, and freedom to exchange in capital and financial markets.

Democracy and constraints to the executive are taken from the Polity IV Project (Center for International Development and Conflict Management CIDCM, University of Maryland) and are available electronically at www.cidcm.umd.edu. Democracy is an additive eleven-point scale (0 to 10) indicator that depends on: competitiveness and openness of executive recruitment, competitiveness of political participation, and constraint on chief executive.

using the full sample of countries, the index of economic freedom enters positively and statistically significantly only for the GMM systems estimators, but its inclusion does not affect our contingency results. Democracy and constraints to the executive enter positively but generally not significantly, again leaving our contingency results intact.

Second, we use export duties and nontariff barriers as alternative trade-barrier measures to test whether the contingency established for tariffs also holds for them. As is well known, both import and export duties introduce an anti-export bias. Since, as reviewed by Edwards (1993), the first generation of cross-country studies on trade and growth focused on the relationship between export growth and GDP growth, it is interesting to test whether export duties exhibit the same contingent relationship that we document for tariffs. Using panel data on export duties from the World Bank, we find no evidence of a relationship between export duties and growth that is contingent on the level of income. This holds across all estimators. In contrast, using cross-sectional OLS we find some evidence of a contingency for nontariff barriers, of the same sort that we document for tariffs. This evidence should be taken cautiously though, because cross-country panel data for a comprehensive measure of nontariff barriers is unavailable. To measure nontariff barriers we use the own-weighted nontariff frequency on intermediate inputs and capital goods obtained from Barro and Lee's (1994) data set.[21] Since this measure is an average between 1985 and 1988, we are only able to test contingency for the 1985-1995 decade. We find that the interaction term between our nontariff measure and the World Bank income rankings is negative and statistically significant, with a p-value of 7%.[22] This suggests that the marginal impact of nontariff barriers on growth is declining in income. More research on this issue would be warranted as more comprehensive data on nontariff barriers becomes available.

Finally, we explore whether the contingent relationship between tariffs and growth on income also extends to other variables. Moreover, we test whether the contingency on income is robust to the allowance for other potential contingencies. For this purpose, we perform two sets of exercises,

[21] This measure includes licensing, prohibitions and quotas. Notice that frequency corresponds to the percentage of products subject to the barrier, but does not capture the degree of restrictiveness.

[22] The sample includes 44 countries, and we control for the same variables of our benchmark estimation (i.e., life expectancy, female and male schooling, and investment and government expenditures as a percentage of GDP).

differing due to data availability. First, we use all four estimation methods as in Table 3, and include two additional interaction terms in the regression specification: the product of tariffs and democracy, and the product of tariffs and population. Democracy is measured using the democracy index variable from Polity IV; log population data were obtained from the World Bank. The motivation behind the consideration of these additional measures is as follows. As noted by Nunn and Trefler (2004), in more democratic countries, governmental preferences tend to be more closely aligned with the welfare of the citizenry. Thus one should expect that in democratic countries, trade protection will be more likely to enhance growth than to reflect rent-seeking and corrupt behavior. If this is the case, then the marginal impact of tariffs on growth should be increasing in democracy (and more generally, in variables measuring the quality of institutions). In addition, if the size of the domestic market (proxied by population) is large, protection may allow local companies to exploit increasing returns, as in Grossman and Helpman (1991) and Rivera-Batiz and Romer (1991). This would imply a marginal impact of tariffs on growth that is increasing in domestic market size.[23]

The upshot of this exercise is that the contingency results reported above are robust to the addition of these two interaction terms. As shown in Table 9 for the GMM systems estimator, the additional interaction terms are insignificant (similar results are found for all other estimators). And although the coefficient on tariffs is no longer universally significant across the different estimators, the interaction terms between tariffs and income levels, and tariff and income rankings, remain significant.[24]

The second exercise is more limited in scope due to data availability. In this case, we are only able to estimate a cross-section regression via OLS for the decade of the 1990s. We do so using World Bank data on paved roads, measured in kilometers relative to total country area as a proxy of infrastructure. If better infrastructure implies that businesses operate in a better environment, then the marginal impact of protection on growth should be increasing in infrastructure. In addition,

[23] Although it is known that the empirical evidence on scale effects in growth rates is weak, the spirit of our exercise is to check the robustness of the contingency we document to other potential contingencies.

[24] Notice that the coefficient on tariffs is just the intercept of the function that relates tariffs and growth.

using UNCTAD data, the World Bank reports data for a few years in the 1990s on average tariffs on primary (food, beverages, tobacco, minerals, nonferrous metals, oil) and manufactured goods (chemicals, manufactured goods of different materials, machinery). We use the ratio of tariffs on primary to manufactured goods as a measure of distribution of tariffs across goods.[25] If the manufacturing sector is more growth promoting than the primary goods sector, as suggested by Grossman and Helpman (1991) and Matsuyama (1992), then the marginal impact of tariffs on growth should be decreasing in the ratio of tariffs on primary to manufactured goods. For instance, Nunn and Trefler (2004) find evidence that countries protecting low-skilled manufacturing grow slower.

In this second exercise, we construct two interaction terms for the decade of the 1990s: the product of tariffs and infrastructure, and the product of tariffs and the ratio of tariffs on primary to manufactured goods. We find (not reported in tables) that both interaction terms are statistically significant (with p-values no greater than 15%) for the 1990s: the marginal impact of tariffs on growth is increasing in infrastructure, while it is decreasing in the protection of primary relative to manufactured goods.[26] However, when all other interaction terms are dropped, the only result that survives is the latter. These results must be read with caution since their scope is limited to the 1990s and to fewer countries than our benchmark exercises. More research is warranted on these issues, particularly on the distribution of tariff barriers across different types of goods. In summary, our finding that the relationship between tariffs and growth is contingent on the level of income is robust to the addition of other potential contingencies.

4 Conclusion

We have examined the relationship between tariff and growth rates from 1975 to 2000, paying particular attention to its potential contingency on the level of economic development. Our analysis

[25] For each country, we use data for the closest year to 1990. Our sample consists of 40 countries.

[26] Regressions include five interaction terms which correspond to the product of tariffs and: income (level or ranking), democracy, log population, infrastructure and relative primary to manufactured tariffs. Results are similar when the interaction terms with democracy and log population are excluded.

has revealed the presence of a significant interaction effect under which the marginal impact of tariffs on growth is declining in initial income. In particular, the relationship between tariffs and growth is negative and significant among the world's rich countries, while positive (with significance exhibiting sensitivity to the particular estimator being employed) among the world's poor countries.

Our results run contrary to the view that higher tariffs are universally detrimental for growth. This is important from a policy perspective, since it indicates that the maintenance of high tariff barriers does not appear to be a leading culprit for the economic stagnation suffered by the world's poor countries.

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Figure 1. Partial Regression Plots for OLS

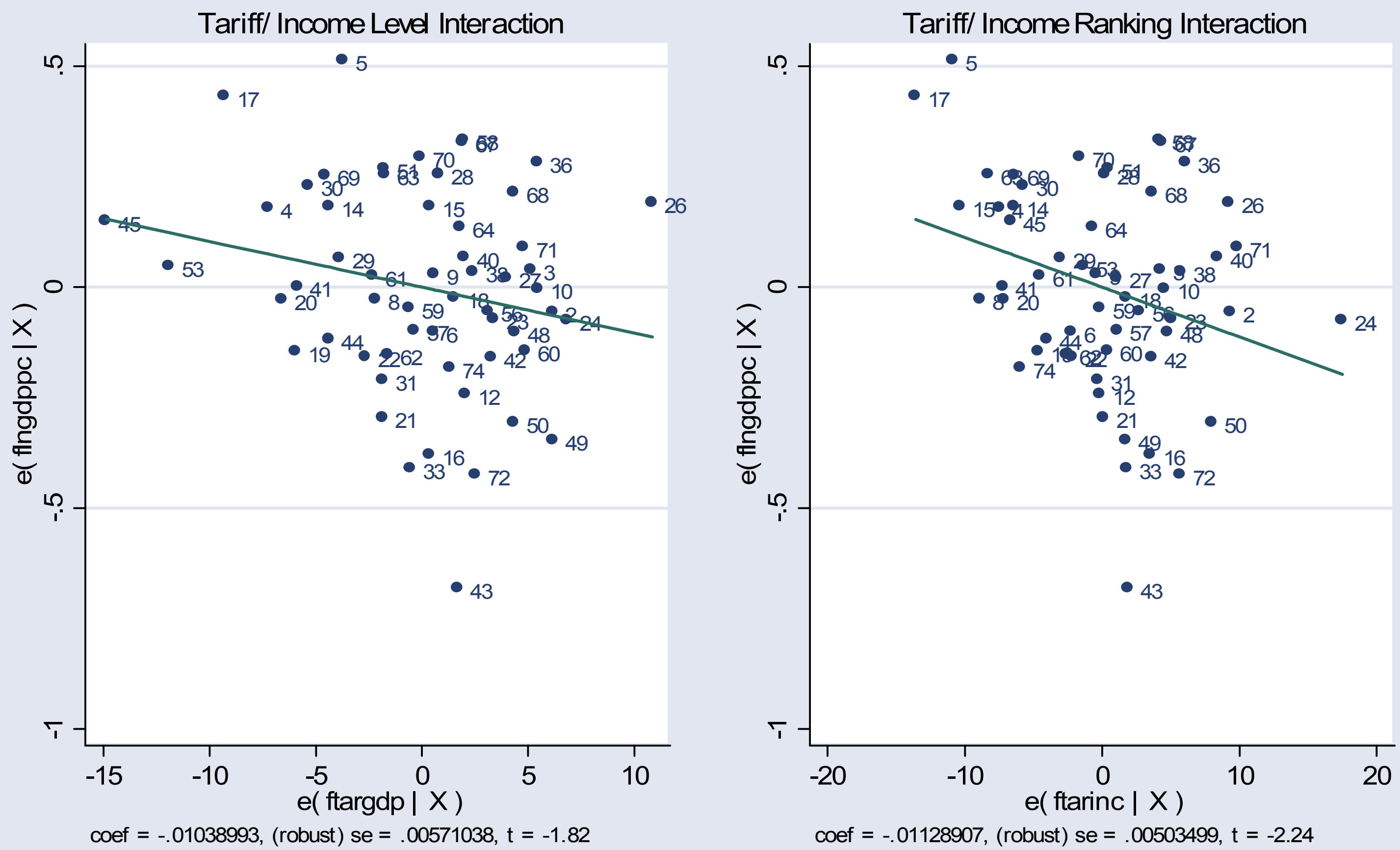

Figure 2. Partial Regression Plots for OLS - Income Sample Splits

Table 1: Summary Statistics

Sample Split	Variable	Observations	Mean	Std. Dev.	Minimum	Maximum
Full Sample	Tariffs	60	10.090	7.846	0.008	33.848
	Growth	60	1.819	1.756	-4.332	6.550
High Income	Tariffs	17	2.687	3.635	0.008	13.926
	Growth	17	1.963	0.633	0.857	3.490
Upper-Mid.	Tariffs	11	5.913	3.426	0.763	12.637
	Growth	11	1.740	1.697	-1.001	5.277
Lower-Mid.	Tariffs	16	12.982	4.812	7.299	22.515
	Growth	16	1.459	2.443	-4.332	6.550
Low	Tariffs	16	17.937	6.895	4.164	33.848
	Growth	16	2.081	1.890	-2.329	4.881
High/ Upper-Mid.	Tariffs	28	3.954	3.841	0.008	13.926
	Growth	28	1.875	1.147	-1.001	5.277
Lower-Mid./ Low	Tariffs	32	15.459	6.367	4.164	33.848
	Growth	32	1.770	2.172	-4.332	6.550

Notes: Growth is measured as annualized growth in per capita GDP; tariffs are measured as import duties as percentages of total imports. See the Appendix for details regarding income definitions.

Table 2: Interaction Specifications, Full-Sample Estimates

2a: Tariff/Income Level Interaction

	OLS			SUR Estimator			GMM Difference Estimator			GMM Systems Estimator		
Regressor	coeff.	s.e.	p value	coeff.	s.e.	p value	coeff.	s.e.	p value	coeff.	s.e.	p value
y_0	0.611	0.097	0.000	0.918	0.027	0.000	0.646	0.136	0.000	0.935	0.047	0.000
tariffs	**0.059**	**0.036**	**0.108**	**0.013**	**0.013**	**0.318**	**0.025**	**0.027**	**0.354**	**0.049**	**0.026**	**0.062**
tar. x y_0	**-0.006**	**0.004**	**0.169**	**-0.002**	**0.002**	**0.302**	**-0.003**	**0.003**	**0.324**	**-0.006**	**0.003**	**0.072**
ln(life)	1.751	0.371	0.000	0.440	0.155	0.006	1.083	0.427	0.014	0.663	0.173	0.000
f.sch.	0.107	0.109	0.328	-0.002	0.033	0.953	0.273	0.117	0.023	-0.006	0.048	0.902
m.sch.	-0.047	0.089	0.601	0.008	0.033	0.812	-0.273	0.118	0.024	-0.010	0.047	0.825
I/Y	0.029	0.007	0.000	0.004	0.002	0.028	0.006	0.004	0.178	0.008	0.002	0.003
G/Y	-0.007	0.006	0.193	-0.002	0.001	0.180	-0.003	0.003	0.174	-0.004	0.002	0.024
Implied λ	0.0200	0.0063	0.0031	0.0172	0.001	0.000	0.0870	0.042	0.042	0.0134	0.010	0.191
	R^2: 0.933			R^2: minimum 0.982 across decades			p values for m1, m2, H tests.: 0.002, 0.707, 0.525			p values for m1, m2, H tests.: 0.000, 0.740, 0.997		

2b: Tariff/Income Ranking Interaction

	OLS			SUR Estimator			GMM Difference Estimator			GMM Systems Estimator		
Regressor	coeff.	s.e.	p value	coeff.	s.e.	p value	coeff.	s.e.	p value	coeff.	s.e.	p value
y_0	0.630	0.098	0.000	0.936	0.025	0.000	0.625	0.135	0.000	0.937	0.049	0.000
tariffs	**0.025**	**0.011**	**0.029**	**0.005**	**0.003**	**0.069**	**0.010**	**0.008**	**0.216**	**0.012**	**0.004**	**0.005**
tar.x rnk	**-0.009**	**0.004**	**0.036**	**-0.003**	**0.001**	**0.030**	**-0.006**	**0.004**	**0.154**	**-0.005**	**0.002**	**0.024**
ln(life)	1.791	0.389	0.000	0.403	0.141	0.005	1.045	0.473	0.031	0.602	0.161	0.000
f.sch.	0.116	0.107	0.282	0.000	0.032	0.998	0.259	0.117	0.031	-0.009	0.059	0.875
m.sch.	-0.051	0.088	0.566	0.005	0.031	0.884	-0.271	0.104	0.012	-0.006	0.054	0.910
I/Y	0.028	0.007	0.000	0.004	0.002	0.040	0.006	0.004	0.098	0.007	0.002	0.003
G/Y	-0.007	0.005	0.183	-0.001	0.001	0.258	-0.003	0.002	0.191	-0.004	0.001	0.005
Implied λ	0.0185	0.006	0.005	0.0131	0.001	0.000	0.0939	0.043	0.035	0.0131	0.010	0.216
	R^2: 0.935			R^2: minimum 0.984 across decades			p values for m1, m2, H tests.: 0.003, 0.942, 0.542			p values for m1, m2, H tests.: 0.000, 0.923, 0.993		

Notes: Estimates are based on 60, 260, 200, and 260 observations, respectively. S.e. denotes heteroskedasticity robust standard error. Constants were estimated but not reported. "Implied λ" is the annualized rate of convergence implied by the coefficient estimate associated with y_0. The m_1 (m_2) test is of the null hypothesis of no first- (second-) order residual serial correlation; H is Hansen's J test of the over-identifying restrictions implied by our choice of instruments. Under our identifying restrictions, the null hypothesis should be rejected under the m_1 test, but not under the m_2 and H tests.

Table 3: Interaction Specifications, Outliers Removed

3a: Tariff/Income Level Interaction

	OLS			SUR Estimator			GMM Difference Estimator			GMM Systems Estimator		
Regressor	coeff.	s.e.	p value	coeff.	s.e.	p value	coeff.	s.e.	p value	coeff.	s.e.	p value
y_0	0.637	0.113	0.000	0.922	0.027	0.000	0.579	0.150	0.000	0.952	0.048	0.000
tariffs	**0.091**	**0.052**	**0.089**	**0.029**	**0.016**	**0.076**	**0.043**	**0.033**	**0.195**	**0.067**	**0.028**	**0.022**
tar. x y_0	**-0.010**	**0.006**	**0.075**	**-0.004**	**0.002**	**0.051**	**-0.006**	**0.004**	**0.183**	**-0.008**	**0.003**	**0.022**
ln(life)	1.767	0.353	0.000	0.406	0.144	0.006	1.175	0.478	0.017	0.601	0.200	0.004
f.sch.	0.167	0.102	0.108	0.023	0.030	0.458	0.315	0.138	0.027	0.032	0.057	0.579
m.sch.	-0.103	0.087	0.241	-0.016	0.031	0.600	-0.328	0.131	0.015	-0.053	0.047	0.262
I/Y	0.030	0.006	0.000	0.005	0.002	0.002	0.005	0.004	0.156	0.008	0.003	0.002
G/Y	-0.007	0.006	0.272	-0.001	0.001	0.270	-0.003	0.002	0.163	-0.003	0.002	0.115
Implied λ	0.018	0.0071	0.0144	0.0162	0.001	0.000	0.109	0.052	0.040	0.0099	0.010	0.328
	R^2: 0.947			R^2: minimum 0.983 across decades			p values for m1, m2, H tests.: 0.008, 0.691, 0.701			p values for m1, m2, H tests.: 0.000, 0.949, 0.999		

3b: Tariff/Income Ranking Interaction

	OLS			SUR Estimator			GMM Difference Estimator			GMM Systems Estimator		
Regressor	coeff.	s.e.	p value	coeff.	s.e.	P value	coeff.	s.e.	p value	coeff.	s.e.	p value
y_0	0.630	0.105	0.000	0.920	0.024	0.000	0.557	0.136	0.000	0.915	0.059	0.000
tariffs	**0.027**	**0.015**	**0.090**	**0.003**	**0.003**	**0.341**	**0.011**	**0.007**	**0.128**	**0.006**	**0.006**	**0.295**
tar.x rnk	**-0.011**	**0.005**	**0.030**	**-0.003**	**0.001**	**0.045**	**-0.007**	**0.004**	**0.057**	**-0.003**	**0.002**	**0.154**
ln(life)	1.815	0.368	0.000	0.325	0.129	0.013	0.959	0.484	0.052	0.425	0.165	0.013
f.sch.	0.180	0.098	0.073	0.019	0.029	0.512	0.295	0.136	0.034	0.036	0.051	0.479
m.sch.	-0.102	0.083	0.225	-0.007	0.028	0.809	-0.306	0.134	0.026	-0.040	0.046	0.389
I/Y	0.028	0.006	0.000	0.005	0.002	0.002	0.004	0.004	0.280	0.007	0.002	0.001
G/Y	-0.007	0.006	0.209	-0.001	0.001	0.337	-0.003	0.003	0.246	-0.004	0.002	0.011
Implied λ	0.0185	0.007	0.008	0.0168	0.001	0.000	0.1172	0.049	0.020	0.0178	0.013	0.175
	R^2: 0.949			R^2: minimum 0.983 across decades			p values for m1, m2, H tests.: 0.004, 0.819, 0.643			p values for m1, m2, H tests.: 0.000, 0.954, 1.000		

Notes: Estimates are based on 56, 242, 186, and 242 observations, respectively. The excluded countries are India, Korea, Mauritius, and Papua New Guinea. Remaining notes are as in Table 2.

Table 4: Measures of Quantitative Significance, Full-Sample Estimates

4a: Tariff/Income Level Interaction

Income Group	OLS			SUR Estimator			GMM Difference Estimator			GMM Systems Estimator		
	impact	s.e.	p value	impact	s.e.	p value	impact	s.e.	p value	impact	s.e.	p value
1	0.525	0.296	0.083	0.153	0.481	0.752	0.181	0.755	0.811	1.309	0.639	0.046
2	0.290	0.287	0.318	-0.116	0.394	0.770	-0.353	0.594	0.555	0.366	0.397	0.361
3	0.104	0.344	0.764	-0.360	0.482	0.458	-0.839	0.811	0.305	-0.492	0.624	0.434
4	-0.040	0.414	0.923	-0.557	0.628	0.379	-1.230	1.117	0.276	-1.182	0.945	0.217

4b: Tariff/Income Ranking Interaction

Income Group	OLS			SUR Estimator			GMM Difference Estimator			GMM Systems Estimator		
	impact	s.e.	p value	impact	s.e.	p value	impact	s.e.	p value	impact	s.e.	p value
1	0.649	0.335	0.058	0.457	0.434	0.298	0.740	0.811	0.366	1.370	0.505	0.009
2	0.292	0.281	0.304	-0.126	0.373	0.736	-0.521	0.525	0.325	0.382	0.466	0.417
3	-0.066	0.318	0.835	-0.710	0.476	0.143	-1.783	1.190	0.140	-0.607	0.736	0.414
4	-0.424	0.423	0.320	-1.293	0.672	0.060	-3.045	2.021	0.138	-1.595	1.109	0.157

Notes: Using the coefficient estimates for tariff and tariff/income interaction terms reported in Table 2, this table reports the differential impact on growth (measured in annual percentage terms) of a 10-percentage-point increase in tariff rates across Income Group rankings. Income Group is 1 for Low-Income Countries; 2 for Lower-Middle Income Countries; 3 for Upper-Middle Income Countries; and 4 for High-Income Countries. Thus the entry under OLS for Group 1 in Table 4a indicates that the estimated impact on average annual growth (in %) of a 10-percentage-point increase in tariff rates for countries in Income Group 1 is 0.525%, etc. Reported p values are associated with t statistics of the null hypothesis that the associated impacts are zero. Impact estimates reported in Table 4a under OLS were computed using as income-interaction terms average values of the natural log of income measured in 1975 calculated within Income Group Classifications. These values are 7.340, 8.286, 9.031, and 9.612. Impact estimates reported in Table 4a under the additional (panel) estimators were computed using as income-interaction terms average values of the natural log of income measured across all "initial" income dates (1975, 1980, etc.) calculated within Income Group Classifications. These values are 7.581, 8.426, 9.194 and 9.812.

Table 5: Measures of Quantitative Significance, Outliers Removed

5a: Tariff/Income Level Interaction

Income Group	OLS			SUR Estimator			GMM Difference Estimator			GMM Systems Estimator		
	impact	s.e.	p value	impact	s.e.	p value	impact	s.e.	p value	impact	s.e.	p value
1	0.567	0.490	0.253	0.117	0.623	0.851	0.159	0.687	0.818	1.382	0.705	0.055
2	0.184	0.349	0.602	-0.467	0.439	0.292	-0.702	0.732	0.342	0.150	0.426	0.726
3	-0.126	0.302	0.679	-1.073	0.484	0.031	-1.595	1.208	0.193	-1.128	0.659	0.093
4	-0.367	0.328	0.268	-1.534	0.658	0.024	-2.275	1.659	0.176	-2.102	1.012	0.043

5b: Tariff/Income Ranking Interaction

Income Group	OLS			SUR Estimator			GMM Difference Estimator			GMM Systems Estimator		
	impact	s.e.	p value	impact	s.e.	p value	impact	s.e.	p value	impact	s.e.	p value
1	0.617	0.455	0.181	0.055	0.482	0.910	0.836	0.835	0.321	0.557	0.765	0.470
2	0.166	0.338	0.626	-0.479	0.387	0.221	-0.601	0.560	0.289	-0.089	0.523	0.865
3	-0.286	0.321	0.378	-1.013	0.457	0.031	-2.038	1.011	0.049	-0.736	0.601	0.227
4	-0.737	0.416	0.082	-1.548	0.641	0.019	-3.476	1.680	0.044	-1.382	0.921	0.140

Notes: Using the coefficient estimates for tariff and tariff/income interaction terms reported in Table 3, this table reports the differential impact on growth (measured in annual percentage terms) of a 10-percentage-point increase in tariff rates across Income Group rankings. Excluded countries are as in Table 3: India, Korea, Mauritius, and Papua New Guinea. The exclusion of these countries altered (relative to those in Table 4a) the average initial income values used as income-interaction terms to calculate estimated impacts in Table 5a. For OLS, the altered values are 7.365, 8.287, 9.031, and 9.612; for the panel estimators, these values are 7.604, 8.385, 9.195, and 9.812. Other notes are as in Table 4.

Table 6: Linear Specifications, Income Sample Splits

6a: Low and Lower-Middle Income Sample

	OLS			SUR Estimator			GMM Difference Estimator			GMM Systems Estimator		
Regressor	coeff.	s.e.	p value	coeff.	s.e.	p value	coeff.	s.e.	p value	coeff.	s.e.	p value
y_0	0.447	0.107	0.000	0.809	0.025	0.000	0.695	0.196	0.001	0.911	0.127	0.000
tariffs	**0.014**	**0.011**	**0.226**	**0.003**	**0.002**	**0.164**	**0.000**	**0.003**	**0.874**	**0.004**	**0.003**	**0.190**
ln(life)	2.460	0.613	0.001	1.022	0.150	0.000	0.331	0.708	0.643	0.594	0.527	0.270
f.sch.	-0.377	0.281	0.195	-0.052	0.033	0.127	0.130	0.216	0.551	0.011	0.083	0.895
m.sch.	0.076	0.232	0.747	-0.026	0.033	0.447	-0.059	0.134	0.664	-0.069	0.088	0.443
I/Y	0.039	0.011	0.003	0.006	0.002	0.001	0.009	0.008	0.291	0.007	0.004	0.076
G/Y	-0.015	0.008	0.089	-0.007	0.002	0.002	-0.007	0.006	0.302	-0.006	0.002	0.015
Implied λ	0.0322	0.010	0.002	0.0425	0.001	0.000	0.0728	0.056	0.202	0.0187	0.028	0.505
	R^2: 0.881			R^2: minimum 0.934 across decades			p values for m1, m2, H tests.: 0.033, 0.394, 0.943			p values for m1, m2, H tests.: 0.005, 0.487, 1.000		

6b: Upper-Middle and High Income Sample

	OLS			SUR Estimator			GMM Difference Estimator			GMM Systems Estimator		
Regressor	coeff.	s.e.	p value	coeff.	s.e.	p value	coeff.	s.e.	p value	coeff.	s.e.	p value
y_0	0.252	0.126	0.062	0.884	0.037	0.000	0.178	0.260	0.500	0.809	0.091	0.000
tariffs	**-0.030**	**0.010**	**0.007**	**-0.011**	**0.002**	**0.000**	**-0.011**	**0.006**	**0.062**	**-0.011**	**0.005**	**0.042**
ln(life)	1.403	0.345	0.001	0.267	0.125	0.045	2.714	1.753	0.134	0.155	0.460	0.738
f.sch.	0.277	0.081	0.003	0.068	0.027	0.020	0.115	0.130	0.385	0.106	0.054	0.062
m.sch.	-0.120	0.048	0.023	-0.041	0.028	0.153	-0.089	0.117	0.455	-0.074	0.056	0.196
I/Y	0.013	0.004	0.002	0.002	0.002	0.331	0.004	0.006	0.537	0.008	0.004	0.043
G/Y	-0.004	0.003	0.204	0.000	0.001	0.846	-0.008	0.005	0.099	-0.001	0.002	0.518
Implied λ	0.0552	0.020	0.008	0.0248	0.002	0.000	0.3450	0.292	0.243	0.0425	0.022	0.065
	R^2: 0.928			R^2: minimum 0.944 across decades			p values for m1, m2, H tests.: 0.280, 0.405, 0.993			p values for m1, m2, H tests.: 0.007, 0.357, 1.000		

Notes: Low and Lower-Middle (Upper-Middle and High) Income OLS estimates are based on 28, 119, 91, and 119 (26, 114, 88, 114) observations, respectively. Relative to the full sample of countries, excluded countries from the low-income group are India, Korea, Mauritius, and Papua New Guinea; excluded countries from the high-income group are Iceland and Venezuela. Remaining notes are as in Table 2.

Table 7: Upward Movers and Asian Tigers

Country	Income Group, 1975	Income Group, 1995	Average Tariff Rate	Average Growth Rate	Coeff.	S.E.	p value
			Full Sample				
						$\boldsymbol{tariff \times y_0}$	
					-0.006	0.003	0.072
			9.612	1.819		$\boldsymbol{tariff \times rank}$	
					-0.005	0.002	0.024
			Remove Outliers				
India	1	1	33.85	3.27		$\boldsymbol{tariff \times y_0}$	
Korea	2	4	7.27	6.54	-0.008	0.003	0.022
Mauritius	2	3	6.50	4.42		$\boldsymbol{tariff \times rank}$	
Papau N.Guniea	2	2	12.26	-0.03	-0.003	0.002	0.154
			Remove Escapees from Low Income Group				
Botswana	1	3	19.19	4.88			
Dom. Republic	1	2	17.40	2.86		$\boldsymbol{tariff \times y_0}$	
Egypt	1	2	25.61	3.79	-0.006	0.003	0.088
Indonesia	1	2	4.16	3.27		$\boldsymbol{tariff \times rank}$	
Sri Lanka	1	2	12.22	2.75	-0.004	0.002	0.054
Syria	1	2	14.61	2.15			
Thailand	1	3	9.72	4.77			
			Remove Escapees from Low / Middle-Income Sub-Sample				
Botswana	1	3	19.19	4.88			
Brazil	2	3	8.36	0.39		$\boldsymbol{tariff \times y_0}$	
Korea	2	4	7.27	6.54	-0.007	0.002	0.005
Malaysia	2	3	6.50	4.42		$\boldsymbol{tariff \times rank}$	
Mauritius	2	3	15.43	4.36	-0.005	0.002	0.054
Thailand	1	3	9.72	4.77			
Turkey	2	3	9.13	1.81			
			Remove Asian Tigers				
Indonesia	1	2	4.16	3.27			
Korea	2	4	7.27	6.54		$\boldsymbol{tariff \times y_0}$	
Malaysia	2	3	6.50	4.42	-0.006	0.004	0.095
Singapore	3	4	0.76	5.28		$\boldsymbol{tariff \times rank}$	
Thailand	1	3	9.72	4.77	-0.004	0.002	0.129

Notes: All estimates obtained using the GMM systems estimator.

Table 8: Low-Tariff Rich and High-Tariff Poor

Country	Income Group, 1975	Income Group, 1995	Average Tariff Rate	Average Growth Rate	Coeff.	S.E.	p value
			Full Sample				
						***tariff* × y_0**	
					-0.006	0.003	0.072
			9.612	1.819		***tariff* × *rank***	
					-0.005	0.002	0.024
			Remove Outliers				
India	1	1	33.85	3.27		***tariff* × y_0**	
Korea	2	4	7.27	6.54	-0.008	0.003	0.022
Mauritius	2	3	6.50	4.42		***tariff* × *rank***	
Papau N.Guniea	2	2	12.26	-0.03	-0.003	0.002	0.154
			Remove Low-Tariff, High Income Countries				
Netherlands	4	4	0.008	1.73		***tariff* × y_0**	
Belgium	4	4	0.011	1.99	-0.006	0.003	0.087
France	4	4	0.060	1.82		***tariff* ×*rank***	
UK	4	4	0.112	2.09	-0.005	0.002	0.029
			Remove High-Tariff, Low Income Countries				
India	1	1	33.85	3.27		***tariff* × y_0**	
Egypt	1	2	25.61	3.79	-0.006	0.006	0.152
Pakistan	1	1	24.25	2.71		***tariff* × *rank***	
Cameroon	1	1	22.39	0.29	-0.004	0.002	0.060
			Remove Both Subsets				
Netherlands	4	4	0.008	1.73			
Belgium	4	4	0.011	1.99		***tariff* × y_0**	
France	4	4	0.060	1.82	-0.007	0.005	0.145
UK	4	4	0.112	2.09		***tariff* × *rank***	
India	1	1	33.85	3.27	-0.004	0.003	0.188
Egypt	1	2	25.61	3.79			
Pakistan	1	1	24.25	2.71			
Cameroon	1	1	22.39	0.29			

Notes: All estimates obtained using the GMM systems estimator.

Table 9: Contingencies on Institutions and Domestic Market Size

9a: Tariff/Income Level Interaction

Regressor	coeff.	s.e.	p value
y_0	0.935	0.038	0.000
tariffs	**0.071**	**0.038**	**0.065**
tar.x y_0	**-0.007**	**0.003**	**0.010**
democracy	**0.003**	**0.003**	**0.269**
tar. x dem	**-0.0003**	**0.0002**	**0.196**
ln(pop)	**0.009**	**0.019**	**0.612**
tar. x pop	**-0.0007**	**0.001**	**0.602**
ln(life)	0.629	0.171	0.001
f.sch.	0.017	0.058	0.765
m.sch.	-0.028	0.057	0.623
I/Y	0.006	0.002	0.012
G/Y	-0.003	0.002	0.095
	p values for m1, m2, H tests.: 0.000, 0.569, 1.000		

9b: Tariff/Income Ranking Interaction

Regressor	coeff.	s.e.	p value
y_0	0.893	0.036	0.000
tariffs	**0.027**	**0.034**	**0.430**
tar.x rnk	**-0.005**	**0.002**	**0.004**
democracy	**0.002**	**0.003**	**0.519**
tar. x dem	**-0.0002**	**0.0002**	**0.396**
ln(pop)	**0.021**	**0.019**	**0.305**
tar. x pop	**-0.001**	**0.002**	**0.578**
ln(life)	0.532	0.157	0.001
f.sch.	0.029	0.056	0.605
m.sch.	-0.022	0.061	0.714
I/Y	0.009	0.004	0.014
G/Y	-0.002	0.002	0.236
	p values for m1, m2, H tests.: 0.000, 0.801, 1.000		

Notes: All estimates obtained using the GMM systems estimator.