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Endogenous trade policy through majority voting: an empirical investigation

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Abstract

The median-voter approach to trade policy determination (within a Heckscher-Ohlin framework) as in Mayer [Am. Econ. Rev. 74(5) (1984) 970] predicts that an increase in inequality, holding constant the economy's overall relative endowments, raises trade barriers in capital-abundant economies and lowers them in capital-scarce economies. We find support for this prediction using cross-country data on inequality, capital-abundance and diverse measures of protection. We perform certain robustness checks that include controlling for the effects of political rights and schooling as well as using alternative datasets on factor endowments.

Keywords: Protection; Openness; Median voter; Inequality

1. Introduction

The median-voter approach, with its focus on majoritarian electoral politics, has been applied quite extensively to diverse political economy issues. This approach and its predictions are best interpreted when the concept of the median voter is not taken literally, but viewed as a convenient analytical device that resolves the conflicting redistributive forces in an unequal society. As Alesina and Rodrik (1994) write:

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> ''We appeal to this (median voter) theorem to capture the basic idea that any government is likely to be responsive to the wishes of the majority when key distributional issues are at stake. Even a dictator cannot completely ignore social demands for fear of being overthrown. Thus, even in a dictatorship, distributional issues affecting the majority of the population will influence policy outcomes.''

In this paper, our aim is to empirically investigate the predictions of the median voter analysis in an important arena of economic policy, namely international trade.[1] Mayer (1984) applies the median voter approach to trade policy determination in standard Heckscher–Ohlin and specific factors (Jones–Ricardo–Viner) trade models. In a two-sector, two-factor (capital and labor) Heckscher–Ohlin version of the Mayer model, the political-economy equilibrium trade policies in an unequal society (one in which the relative capital endowment of the median individual is less than the mean) will be biased in favor of labor. More trade results in a higher factor reward for the abundant factor and a lower factor reward for the scarce factor. Hence, the model predicts an equilibrium trade policy biased against trade in capital-rich countries and in favor of trade in capital-scarce economies. However, in the real world, trade policies are almost everywhere and always biased against trade. This discrepancy between the median-voter prediction and the empirical evidence can be attributed to other kinds of redistributive pressures on the government, such as those from lobbies and special-interest groups. Real-world politics consists of both special interest and majoritarian politics among numerous other things. Thus, there might exist in reality elements of redistributive pressures similar to those captured by the median-voter argument but may be rendered invisible by other opposing elements.

Our paper, therefore, focuses on a second important prediction that can be derived from this median-voter approach to trade policy determination. This prediction is about cross-country variations in levels of trade barriers and *not* about the actual orientations (signs) of the levels. More precisely, we perform a simple comparative-static exercise in the Mayer–Heckscher–Ohlin (Mayer-H-O) framework to obtain the result that an increase in inequality (the difference between the mean and the median capital-labor ratio), holding constant the economy's overall relative endowments, raises trade barriers in capital-abundant economies and lowers them in capital-scarce economies.[2] It is exactly this

[1] The other approach focusses on pressure group politics. See, for instance, Feenstra and Bhagwati (1982), Findlay and Wellisz (1982), Grossman and Helpman (1994), Hillman (1989), Magee et al. (1989) and Mitra (1999).

[2] An increase in inequality increases the demand for redistribution from capital to labor. This can be achieved through trade policies that increase further the factor reward to labor but reduce the reward to capital, which in turn is achieved by increasing the domestic price of the labor-intensive good in a two-sector, two-factor Heckscher–Ohlin economy. Thus, an increase in inequality would result in a tightening of trade restrictions in capital-abundant economies and their reduction in capital-scarce countries.

prediction about cross-country variation in trade policy that we are able to investigate empirically using cross-sectional data on inequality, capital-abundance and diverse measures of trade restrictions and openness.

It is important to note here that in the Mayer-H-O framework, an increase in inequality makes the import tariff more positive (i.e. makes trade policies more antitrade) in a capital-abundant economy, while as a result of such an increase in inequality, the import tariff becomes more and more negative (i.e. trade policies become more protrade) in a labor abundant economy. However, as argued above, in the real world there are possibly other components of the tariff (arising from other factors or considerations) which are, in combination, always positive enough to make the overall tariff positive in countries of all degrees of capital abundance or scarcity. *Holding these other effects constant with respect to inequality, the overall import tariff can rise or fall with inequality to the extent that the positive or negative Mayer component becomes more positive or more negative*.

An alternative political economy model using a lobbying approach in a static setting (within the same two-sector, two-factor Heckscher–Ohlin framework) makes exactly the opposite prediction. As asset or capital inequality increases, the ownership of capital becomes more concentrated. Thus, in this one-period setting, we have a reduction in the free-riding related public-good provision problem for pro-capital lobbying relative to pro-labor lobbying. The result is an intensification of redistribution from labor to capital (as opposed to redistribution from capital to labor in the median-voter case). Thus, in a single-period lobbying model with capital and labor pitted against each other in a two-factor, two-sector Heckscher–Ohlin environment, an increase in inequality will lead to an increase in import protection in capital-scarce economies, but will lead to a reduction in import-protection in capital abundant economies.[3]

In a repeated-game setting, Pecorino (1998) shows that the effect of concentration on the free-rider problem is ambiguous. A reduction in concentration increases the current-period gain from defecting from the cooperative outcome and at the same time lowers the profits that each firm makes when punished (through a non-cooperative regime) as a result.[4] Thus, what we learn from the above lobbying

[3]The simplest way to obtain such a result is to consider a model of the type presented in Rodrik (1986) and perform a comparative static exercise there by varying the the number of capitalists, holding constant the aggregate stock of capital and the total population. In such a model, pro-capital redistribution is increasing in asset inequality, while in the median-voter model, pro-labor redistribution is increasing in asset inequality.

[4]Magee (2001) uses a political-contributions approach (and endogenizes the tariff-formation function used by Pecorino) in a repeated game setting to resolve some of this ambiguity. He is able to characterize the equilibria under different sets of parameter values and is able to derive conditions on the parameters (such as the discount rate and the government's bargaining power with respect to the import-competing lobby) that lead to monotonically increasing (decreasing) or even non-monotonic relationships between concentration, the ease of free riding and the maximum sustainable tariff. Which of these parameter values real-world lobbies face is not easy to identify.

models is that the relationship between inequality or concentration and trade protection can only be resolved empirically.

We employ two measures of inequality (or equality) in studying its impact on different measures of trade restrictiveness. One is the Gini-coefficient which is a summary measure of inequality and is consistent with a broad interpretation of the model we employ. The other is the share of the median quintile of the population in national income, which fairly accurately corresponds to the share of the median voter in the Mayer framework. In looking at the effects of inequality, we allow it to change direction and magnitude as the relative factor proportions change when we move across countries. We carry out our empirical investigation using three separate measures (Nehru–Dhareshwar, Summers–Heston and Easterly–Levine) of relative factor endowments (capital per worker), constructed using different methods and under different assumptions. Across all measures of trade restrictiveness and using different measures of the capital-labor ratio, we find strong evidence in favor of the above-mentioned median voter prediction. An increase in the Gini-coefficient or a reduction in the median quintile's share, holding constant the economy's overall relative endowments, does in fact, raise trade barriers in capital-abundant economies and lowers them in capital-scarce economies.[5] Further, this result is extremely robust to the use of controls.

In this context, our result is consistent with the results of econometric studies (using micro-level suvey data) on individual level trade policy preferences such as Balistreri (1997), Beaulieu (2001), and Scheve and Slaughter (1998). These authors find that for both Canada and the US in recent years, factor type has been the dominant determinant of support for or opposition to trade barriers. Individuals owning proportionally more of the scarce factors are in favor of trade barriers, while those owning proportionally more of the abundant factors do not like trade restrictions. If we can take the empirical findings on individual trade policy preferences as given, a simple application of the median-voter analysis over these preferences should theoretically deliver our non-monotonic results. For the sake of completeness, we present a framework in which even the empirically observed type of individual preferences over trade policy are derived from first principles using a two-sector, two-factor Heckscher–Ohlin set up.

The main theoretical proposition presented in our paper is driven by the Stolper–Samuelson effect and therefore, our empirical results can be interpreted only in the context of this effect. Besides the above individual-level, revealed preference evidence for the Stolper–Samuelson theorem, there are papers that have found support for it using data on Political Action Committee (PAC) contributions and congressional voting patterns. These studies specifically find support for the two-factor, capital-labor version of the Stolper–Samuelson effect and thus are specially relevant for our empirical investigation. Beaulieu (2000) finds some

[5]In addition to using inequality as a variable, an interaction term between inequality and the capital-labor ratio is used to endogenously determine from the data the threshold capital-labor ratio where the trade restrictiveness-inequality relationship changes sign or direction.

evidence of congressional voting patterns on trade policy in the US being affected by the factor-endowment composition of constituencies. One of the interesting empirical regularities unravelled by his study is a negative relationship between the likelihood that a candidate votes in favor of the CUSTA, GATT or NAFTA and the size of contributions from labor PACs. He also finds a positive effect of contributions from capital (corporate) PACs in the case of the CUSTA. In the case of the GATT and NAFTA, however, he finds no effect of capital contributions. Kahane (1996) finds that after controlling for state characteristics, the likelihood of voting against the NAFTA in both the House and the Senate was increasing in contributions by labor PACs. Steagall and Jennings (1996) find that the likelihood of a favorable House vote for the NAFTA was again decreasing in labor contributions, but also increasing in capital (corporate) contributions. However, contributions are endogenous to political and other leanings of the candidate. Baldwin and Magee (2000), after taking into account this endogeneity, find strong evidence that the likelihood of a favorable vote for NAFTA or GATT cast in the House was decreasing in labor contributions but increasing in business contributions. Beaulieu and Magee (2000) determine industry affiliation of these capital and labor PACs. They find that both the probability of a capital PAC contributing money to a candidate and the size of its contribution to a candidate were higher if he/she was a supporter of NAFTA, while the reverse was true for a labor PAC. Industry affiliation of these PACs did not seem to matter in their contributions decisions in this NAFTA context.[6]

The contribution of our paper is two-fold. Firstly, our results uncover a robust empirical regularity in the relationship between trade protection and inequality and provide some credibility to the median voter approach to political economy. Secondly, the paper adds to the empirical literature on cross-national variation in protection.[7]

[6] In contrast to the studies mentioned above, earlier studies (using older data) by Irwin (1994, 1996) and Magee (1978) find that industry of employment was the major determinant of individual level trade policy preferences in the British elections of the early twentieth century and in the testimonies of trade unions, management and industry associations before the House Ways and Means Committee on the Trade Reform Act of 1973 in the US, respectively. However, Rogowski (1987) shows how coalitions formed in the US, Britain and Germany in the nineteenth century are those predicted by the Heckscher–Ohlin model.

[7] See Rodrik (1995) for a discussion of the importance of (and the need for) empirical work on cross-country variations in protection.To our knowledge, there are only two cross-country empirical studies on protection. Magee et al. (1989) (Chapter 16) find that average tariff rates tend to decrease as capital-labor ratios increase. Mansfield and Busch (1995) examine cross-national variation in average protection levels among 14 advanced industrial countries pooled over 2 years, 1983 and 1986. They find that non-tariff barriers are increasing in country size, unemployment rate and number of parliamentary constituencies and are higher for countries that use proportional representation as their electoral system. Also, there are three well known cross industry studies on protection in the US—Goldberg and Maggi (1999), Gawande and Bandopadhyay (2000) and Trefler (1993)—all of which focus on the predictions of lobbying/political contributions models about cross-industry variation in protection.

In Section 2, we present a modified version of the Mayer (1984) model and perform a comparative-static exercise to derive the implications of increasing inequality for trade policy determination. Section 3 describes the specification of the econometric model and explains the various inferences that the model allows. Section 4 briefly discusses the data and the choice of regressors. In Section 5, we discuss our empirical results and finally, in Section 6, we make some concluding remarks.

2. Theoretical framework

Let us consider a two-factor, two-sector, small-open, Heckscher–Ohlin economy. Good 1 is the importable and good 2 the exportable. The domestic price of the importable is p while its world price is p^*, so that $p = p^*(1 + t)$. Let good 2 be the numeraire good. Both goods require both capital and labor in their production carried out under constant returns to scale. On the demand side, individual preferences are taken to be identical and homothetic. An individual h's indirect utility function can, therefore, be written as $V(p)I^h$.

For simplicity, we assume that each individual owns one unit of labor and k^h units of capital. Let the share of an individual h in the overall capital stock of the economy be denoted by σ^h, so that $k^h = \sigma^h K$ where K is the aggregate capital stock of the economy.[8] Let L be the total number of individuals and hence the aggregate labor endowment of the economy. The income of an individual is then given by

$$I^h(p) = w(p) + r(p)\sigma^h K + \phi^h(p - p^*)M(p) \tag{1}$$

where ϕ^h is the share of an individual h in the total tariff revenue (*equal to an individual's share in factor income by assumption*) and $M(p)$ the imports of good 1. $w(p)$ and $r(p)$ are the wage rate earned by labor and rental on capital, respectively, both being solely the functions of the domestic price of the importable.

An individual h's most preferred tariff is determined by maximizing $V(p)I^h(p)$ with respect to p, which yields

[8] Our assumption, in this theory section, that $L^h = 1$ for all h, is a simplifying assumption. The model can be easily modified to incorporate heterogeneity in labor holdings so that individuals differ in terms of $k^h = K^h/L^h$, the relative capital–labor endowment. In that case, the equilibrium tariff will be determined by the ratio of median K/L to the average K/L. However, whether this makes a difference depends on our interpretations of K and L, i.e. whether capital is interpreted to be just physical capital or also includes human capital and how labor is being measured-in terms of the number of workers or in terms of efficiency units. Our measure of labor (in the data we use) is in terms of the number of workers (i.e. one unit of labor per person) and we interpret the income gini (used in our empirical investigation) as reflecting heterogeneity in both physical and human capital.

$$t^h = \frac{-I}{p^* M'(p)} \frac{\partial \phi^h / \partial p}{\phi^h} \tag{2}$$

where I is aggregate income. Imports are negatively related to the domestic price of the importable and so we have $M'(p) < 0$. Furthermore, $\partial \phi^h / \partial p > (<) 0$ if individual h is relatively well (poorly) endowed in the factor used intensively in the production of the importable and consequently such an individual's most preferred tariff will be positive (negative).

Assuming that the voters differ only along a single dimension, namely in their relative capital–labor endowment k^h and that there are no voting costs, the tariff under majority voting can be obtained using the median voter theorem and is the one that maximizes the utility of the individual with the median relative capital–labor endowment in the economy. In other words, it is obtained by maximizing $V(p)I^{mv}$ where mv stands for the median voter. This is equivalent to maximizing $v(p) + i^{mv}$ where $v(p) = \ln V(p)$ and $i^{mv} = \ln I^{mv}$. It is assumed that this objective function is concave with respect to price. Expanding the expression for i^{mv}, we have $i^{mv} = \ln[w(p) + r(p)\sigma^{mv}K] + \ln[1 + \delta(p, K/L)]$ where δ is the ratio of total tariff revenue to national factor income. The first order condition of our maximization problem gives us

$$v'(p) + \partial i^{mv} / \partial p = 0 \tag{3}$$

σ^{mv} is the median voter's share in the capital stock and is always below the average share in real world distributions (see Alesina and Rodrik, 1994). Thus, σ^{mv} can be considered to be an inverse index of inequality or an index of equality in the distribution of assets. Therefore, in order to study the effect of a change in the degree of inequality in asset distribution on the nature of trade policy, we look at the effect of a change in σ^{mv}, holding constant the economy's aggregate factor endowments. Let t^{mv} be the median voter's most preferred level of tariff (on the importable), also called the political–economic equilibrium tariff. Differentiating our first order condition to perform comparative statics we obtain

$$\frac{\partial t^{mv}}{\partial \sigma^{mv}} = \frac{-[r'(p)w(p) - r(p)w'(p)]K}{p^*[w(p) + r(p)\sigma^{mv}K]^2[v''(p) + \partial^2 i^{mv} / \partial p^2]} \tag{4}$$

Since an increase in the domestic price of the importable increases the reward to the scarce factor and reduces that for the abundant factor, we have $r'(p) < 0$ and $w'(p) > 0$ for a capital-abundant country, while $r'(p) > 0$ and $w'(p) < 0$ for a labor-abundant country. The denominator is always negative due to the restriction of concavity imposed on the objective function. Thus, the above derivative is negative when the economy is capital abundant, so that an increase in inequality leads to an increase in the equilibrium tariff. For a labor abundant country, the above derivative has a positive sign. In other words, an increase in inequality always results in an increase in the demand for redistribution through policies that would benefit labor at the expense of capital. In a capital abundant country, the

importable is the labor intensive good and an increase in the demand for redistribution from capital to labor would represent a demand for policies that increasingly favor the importable sector. In a labor abundant economy, the importable sector is the capital intensive sector and hence more redistribution towards labor requires policies that are more biased against the importable sector. This leads us to the following proposition whose empirical validity we test in this paper.

Proposition. *Holding other things constant, an increase in inequality leads to more restrictive or less open trade policies in capital abundant countries, while it leads to less restrictive or more open trade policies in capital scarce economies.*

While the predictions are not as precise once we allow for more than two factors,[9] we will attempt to argue that the median voter predictions stated in the above proposition are not as specific to the two-factor framework as they appear. First, let us assume that there are three factors-physical capital (K), human capital or skills (H) and raw, unskilled labor (L). Per capita income in any country is $(\text{National Income})/L = r(K/L) + w_H(H/L) + w$ where w_H denotes the return on human capital. Thus, in order to be rich (poor), countries have to be relatively abundant (scarce) in K and H combined or relatively scarce (abundant) in L. In any country, concentration in the ownership of skills and physical assets leads to inequality.[10] Higher inequality (of this kind) implies greater dependence for the majority of the population on their raw, unskilled labor power, and a greater demand for redistribution (see Alesina and Rodrik, 1994), thereby leading to prolabor redistribution policies. In rich (poor) countries, this leads to higher (lower) trade barriers.[11]

[9]The median voter model is usually applied when individuals differ along a single dimension—in this case the capital labor ratio, which when combined with a monotonicity result, yields single peaked preferences. So even when allowing for multiple factors and heterogeneity in their ownership, it is crucial that voters and individuals differ along a single dimension.

[10]Endowments of physical and human capital should be correlated (both at the country and individual levels), as it is the marginal rate of time preference that determines the steady state levels of both in the absence of credit market imperfections, while in the presence of such imperfections, the ownership of physical assets directly affects the ability to acquire skills.

[11]If we go beyond three factors, our basic result qualitatively will still hold though it might be weakened a bit. The higher dimensional version of the Stolper-Samuelson theorem implies that if a factor is 'scarce enough' ('abundant enough'), it will be helped (harmed) by trade barriers (see Leamer and Levinsohn, 1995). Consider a continuum of types of skills (high level or high paying to low level or low paying) and types of physical assets (high tech and high return like computers to low tech and low return like hammers, screw-drivers, etc.). Further it would be realistic to assume that in all countries (rich and poor) the majority will possess the lower end of skills and assets. Rich countries will be abundant in the higher end factors, while poor countries in the lower end factors. Under these conditions, an increase in inequality through higher concentration of higher level assets and skills in the hands of fewer individuals, should increase the demand for redistribution from rich to the poor. Then, at least, in the very rich countries, this will generate high trade barriers and in the very poor ones lead to lower barriers.

3. Econometric methodology

The comparative static result of the previous section provides the foundation for our empirical work. In countries with high (K/L) ratios, inequality and trade restrictiveness should be positively related, but when (K/L) is low there is an inverse relationship between these two variables. A priori, we do not know at what level of (K/L), the relationship changes sign. The following specification takes care of this problem by allowing the data to tell us the exact location of this turning point:

$$\mathrm{TR}_i = \alpha_0 + \alpha_1 \mathrm{INEQ}_i + \alpha_2 \mathrm{INEQ}_i \times (K/L)_i + \alpha_3 (K/L)_i + \mathbf{X}_i \boldsymbol{\beta} + \epsilon_i \tag{5}$$

where TR_i is the extent of trade restrictions in country i, INEQ_i is the level of inequality, $(K/L)_i$ the capital–labor ratio and $\mathbf{X}_i$ is a row vector of control variables.[12,13] Taking the partial derivative of TR_i with respect to INEQ_i, we have

$$\frac{\partial \mathrm{TR}_i}{\partial (\mathrm{INEQ})_i} = \alpha_1 + \alpha_2 (K/L)_i \tag{6}$$

The prediction of the comparative static exercise of the previous section is that $\alpha_1 < 0$ and $\alpha_2 > 0$ such that $\alpha_1 + \alpha_2 (K/L)_i \gtreqless 0$ as $(K/L)_i \gtreqless (K/L)^*$ where $(K/L)^* = -\alpha_1/\alpha_2$ is the turning point capital–labor ratio determined endogenously from the data, given our estimating equation. Another requirement for the prediction to hold is that $(K/L)^*$ should lie within the range of values of (K/L) in the dataset, i.e. $(K/L)^{\mathrm{MIN}} < (K/L)^* < (K/L)^{\mathrm{MAX}}$.

We start with the basic regression in which TR is regressed on (K/L), INEQ and $\mathrm{INEQ} \times (K/L)$. The inclusion of (K/L) as a separate variable (in addition to INEQ and $\mathrm{INEQ} \times (K/L)$) allows $\partial \mathrm{TR}_i / \partial (K/L)_i$ and the variable component of $\partial \mathrm{TR}_i / \partial (\mathrm{INEQ})_i$ to differ in sign. Otherwise, they are restricted to having the same sign. We then add controls such as schooling and democracy to see whether our results are robust to their inclusion. The reasons for their inclusion as controls is explained in Section 5.

[12] In our estimation, we use the capital–labor ratios in natural logs. The reasons are as follows: (1) For all measures (TARIFF, IMPORT DUTY, QUOTA and $(X+M)/\mathrm{GDP}$), we have two to three outliers in regressions that use logs (of K/L), while there are 18–24 outliers in each regression when we use K/L in levels. Moreover, in the case of logs, the results are robust to the deletion of outliers. (2) For all protection measures, the J-test for the model with log (of K/L) vs. the model with the level clearly accepts the former as the null hypothesis against the latter as the alternative and rejects the latter as the null against the former as the alternative.

[13] We have also tested for non-linearities/non-monotononicities with respect to K/L by additionally including its square. This additional term is statistically insignificant at 15% and even much higher levels. Also, squares of INEQ and K/L thrown in simultaneously were statistically extremely insignificant. We also could not detect any non-linearities (at 15% and even higher levels of significance) in any of our variables (K/L, inequality, their cross product and other control variables) when we performed the Ramsey Reset test for all our regressions, both with and without controls. The detailed results are available at www.ualberta.ca/economics/dm/dm.htm.

As mentioned before, one index of inequality we use is the Gini-coefficient which is a broader measure than the interpretation of inequality used in the theoretical framework in Section 2. Alternatively, we use $Q3$, the share of the third quintile in national income, which corresponds much closer to the share of the median voter in the theoretical model. This is an inverse measure of inequality (or rather a measure of equality) and so the signs of the coefficients of this variable and its interaction with K/L are expected to be the reverse of those obtained when the Gini-coefficient is used.

We also do a few other robustness checks. Our measures of inequality are the income Gini-coefficient and the median quintile's share in national income or expenditure, both of which are indirect measures of asset inequality (or equality) since they are actually measures of income inequality.[14] There is the possibility of reverse causation running from trade policy to income inequality. Moreover, in a more dynamic context (for example in a multisectoral Solow model), K/L may be endogenous with respect to trade policy. Protection, by affecting the production structure, can affect accumulation and the steady state level of the capital stock. Even though our right-hand side variables generally are lagged with respect to the ones on the left-hand side, this would not take care of the endogeneity problem in cross-sectional analysis when variables exhibit stickiness. Therefore, we use tests suggested by Hausman (1978) and Smith and Blundell (1986). In a linear model, Hausman (1978) showed that an easy way of implementing the Hausman test for exogeneity is to first run reduced form regressions of each of the variables (in our case, INEQ, K/L and INEQ*K/L) that are suspected to be endogenous on all the exogenous variables from our main regression and other exogenous variables which theory suggests might affect any of these endogenous variables. The second step involves computing the residuals from each of these auxilliary regressions and inserting them as additional right-hand side variables in our main estimating regression. If these residuals are jointly significant (insignificant), our plain OLS estimation of the model produces inconsistent (consistent) estimates. However, in the case of the joint significance of auxilliary residuals and the consequent endogeneity, this Hausman regression will produce coefficient estimates that are consistent and identical to IV estimates. The standard errrors of the coefficient estimates, however, need to be corrected by multiplying those from the Hausman regression by an appropriate correction factor (which we do when required).

[14]The only measure of asset inequality on which cross-country data is available is the land-Gini. However, using data on land-ownership inequality directly in our regressions is not very meaningful, especially in a Heckscher–Ohlin framework. See the appendix for a more detailed analysis of the interpretation of regression coefficient estimates using income inequality as opposed to direct measures of asset inequality.

4. Data sources and some basic statistics

The detailed description of the data and their sources and the dataset itself are available at the following website: www.ualberta.ca/economics/dm/dm.htm. Here, we provide a very brief summary of the data used in this paper. Our dependent variable is trade protection and our independent variables of interest are inequality, the capital–labor ratio, indicators for democracy and political rights, and schooling. For the regional effects, we will be using region-specific dummies.

To test for the robustness of our results, we use a variety of trade policy measures: total import duties collected as a percentage of total imports (IMPORT DUTY), an average tariff rate calculated by weighing each import category by the fraction of world trade in that category (TARIFF)[15], a coverage ratio for non-tariff barriers to trade (QUOTA) and an indirect measure of trade restrictions-the magnitude of trade flows relative to GDP, defined as $(X + M)/\text{GDP}$.

The degree of income inequality is measured by the Gini-coefficient and alternatively, by an inverse index-the median quintile's share in national income or expenditure.

Using the Nehru–Dhareshwar data on capital in conjunction with the data on labor (defined as population between ages 15 and 64), we calculate the capital–labor ratio. The average for the 1980s is used. The data on capital stock at 1987 domestic prices are converted into 1987 constant dollars using the 1987 exchange rate. We perform robustness checks using the Summers–Heston and the Easterly–Levine capital per worker data whose country coverage is much smaller than the Nehru–Dhareshwar data.[16]

For a measure of democracy, we use the Freedom House (Gastil) measure of democracy that provides a subjective classification of countries on a scale of 1 to 7 on political rights, with higher ratings signifying less freedom. Again, the average for the 1980s is used. We use 'schooling' as a control variable, where schooling is defined as the average number of schooling years in total population over the age of 25. The summary statistics for these variables and the correlation across the various measures of trade restrictions are available at www.ualberta.ca/economics/dm/dm.htm.

For the Hausman regressions, the additional variables required are civil liberty (another Gastil index), schooling, $M2/\text{GDP}$, the Gini-coefficient for the distribution of land, savings rate and the population growth rate. The population growth rate and the savings rate are parameters in the Solow growth model in

[15]The variable is referred to as tariffs, although it includes all import charges, such as duties and customs fees.

[16]The Summers–Heston data are in constant 1985 international dollars, i.e. conversion into a common denominator is based on cost differentials and not the law of one price. The Easterly–Levine data are constructed using the Summers–Heston disaggregated sectoral investment data along with information on disaggregated sector-level depreciation, etc. to arrive at more accurate measures.

which the steady state per-capita capital stock is determined endogenously, while Li et al. (1998) explain intertemporal and international variation in income inequality in terms of variations in political and civil liberties (the Gastil indices), schooling, $M2/GDP$, and the Gini-coefficient for the distribution of land.

5. Results

Figs. 1 and 2 show some simple tariff vs. Gini scatter plots for capital-abundant and labor-abundant countries, respectively. The median capital–labor ratio from the Nehru–Dhareshwar dataset is used to classify countries as capital and labor abundant. There is clearly a positive correlation between tariffs and inequality in the case of capital-abundant countries and a negative correlation between the two for labor-abundant countries. Thus, even the very basic methods of data analysis can provide support for the theory presented in this paper.

5.1. OLS regressions (with and without controls)

Tables 1 and 2 present the regression results (with and without controls) for our main estimating equation (Eq. (5) in Section 3 of this paper). The sample size which ranges from 44 to 64, depends on the country coverage of the data on the different variables used. The regression models as a whole are always significant at the 5% level. The R^2 ranges from 0.15 (in the case of quota without controls and

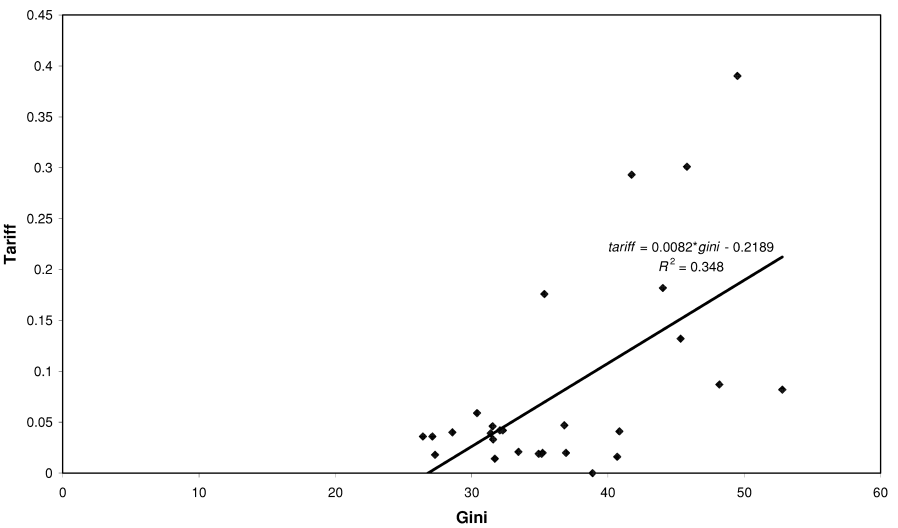

Fig. 1. Relationship between tariff and inequality in capital-abundant countries.

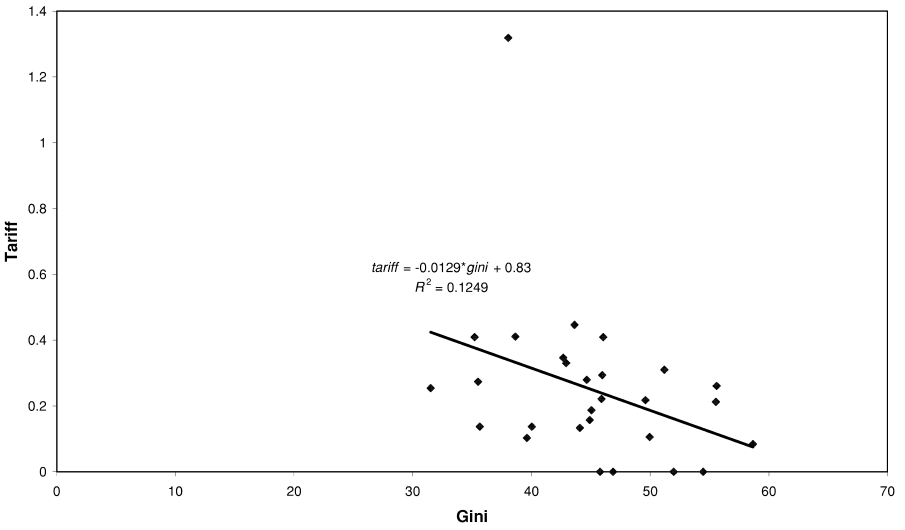

Fig. 2. Relationship between tariff and inequality in labor-abundant countries.

using the median quintile's share $Q3$) to 0.60 (in the case of import duty with controls, using the Gini coefficient).

Apart from the regressions based upon the quota coverage ratio, we find strong support for the predictions of the median voter model, both with and without controls. As predicted, $\alpha_1 < 0$ and $\alpha_2 > 0$ when the Gini coefficient is used and the reverse when $Q3$ is used.[17] For quotas, these coefficients are insignificant. In the quota regressions using the Gini coefficient, they even have the wrong signs.

The quota coverage ratio suffers from measurement error problems due to smuggling, coding problems and weaknesses in the underlying data. It also does not distinguish between highly restrictive barriers and non-binding ones, thus suggesting only their existence and being unable to measure their effect on imports. Harrigan (1993) has found that for OECD countries in 1983 both price and quantity NTB coverage ratios are, in most cases, not associated with lower imports. He points out that these coverage ratios are the noisiest indicators of trade policy as there are severe problems with their construction procedure and are not conceptually what is desired.[18] However, in subsequent sub-sections (where we use alternative capital per work measures), we show results even for the quota variable that are consistent with median-voter predictions.

[17]We also find support for the median-voter predictions using the Sachs–Warner binary measure of openness as the dependent variable in a logistic regression. However, we do not present those results as the Sachs–Warner measure has come under heavy criticism recently.

[18]For a detailed discussion of the problems with quantity and price NTB coverage ratios, see Leamer (1990).

Table 1
Gini coefficient-regression with and without controls

	Tariff	Quota	Import duty	$(X+M)$/GDP	Tariff	Quota	Import duty	$(X+M)$/GDP
Gini	−0.029**	0.012	−1.048*	0.049	−0.051***	0.016	−2.141***	0.082*
	(0.016)	(0.022)	(0.679)	(0.04)	(0.023)	(0.03)	(0.794)	(0.05)
Gini*capital–labor ratio	0.003***	−0.0002	0.144***	−0.006*	0.006***	−0.001	0.249***	−0.01***
	(0.001)	(0.002)	(0.067)	(0.004)	(0.002)	(0.003)	(0.083)	(0.005)
Capital–labor ratio	−0.189***	−0.037	−8.406***	0.341***	−0.328***	−0.03	−13.85***	0.58***
	(0.068)	(0.093)	(2.888)	(0.166)	(0.11)	(0.151)	(3.986)	(0.247)
Schooling					0.006	−0.03	0.593	−0.042
					(0.019)	(0.029)	(0.705)	(0.042)
Political rights (Gastil)					−0.025	−0.02	0.397	0.037
					(0.021)	(0.028)	(0.776)	(0.047)
Sub-saharan Africa					−0.068	−0.259***	1.537	0.022
					(0.095)	(0.126)	(3.341)	(21.7)
East Asia					−0.103	−0.202**	−6.981***	0.591***
					(0.088)	(0.119)	(3.189)	(0.203)
Oil					0.017	−0.107	1.085	−0.141
					(0.081)	(0.108)	(3.052)	(0.187)
Constant	1.942***	0.157	78.218***	−2.372	3.291***	0.414	129.835***	−4.365**
	(0.71)	(0.96)	(29.541)	(1.722)	(1.005)	(1.356)	(35.938)	(2.273)
No. of observations	51	50	50	56	45	45	44	49
R^2	0.36	0.25	0.44	0.22	0.47	0.42	0.6	0.37
F-statistic	8.82***	5.23***	11.97***	5.03***	4.01***	3.29***	6.79***	3.0***
Critical capital–labor ratio	9.7	55	7.3	8.2	8.5	16	8.6	8.2

Standard errors in parantheses; *** significant at 5% level, ** significant at 10% level, * significant at 15% level.

Table 2
Third quintile-regression with and without controls

	Tariff	Quota	Import duty	$(X+M)$/GDP	Tariff	Quota	Import duty	$(X+M)$/GDP
$Q3$	0.136***	0.081	6.752***	−0.196***	0.121***	0.021	5.15***	−0.09
	(0.042)	(0.06)	(1.821)	(0.062)	(0.055)	(0.075)	(1.879)	(0.068)
$Q3$*capital–labor ratio	−0.015***	−0.012**	−0.747***	0.025***	−0.013***	−0.005	−0.558***	0.012**
	(0.004)	(0.006)	(0.183)	(0.062)	(0.006)	(0.008)	(0.2)	(0.007)
Capital–labor ratio	0.18***	0.158**	8.803***	−0.324***	0.15***	0.07	6.508***	−0.154*
	(0.057)	(0.082)	(2.215)	(0.086)	(0.078)	(0.106)	(2.647)	(0.096)
Schooling					−0.025	−0.054***	−1.073**	0.041**
					(0.018)	(0.025)	(0.602)	(0.021)
Political rights (Gastil)					−0.015	−0.045**	−0.602	−0.001
					(0.018)	(0.025)	(0.653)	(0.023)
Sub-saharan Africa					−0.081	−0.162*	1.003	0.023
					(0.083)	(0.112)	(2.673)	(0.097)
East Asia					−0.051	−0.052	−2.717	−0.002
					(0.097)	(0.133)	(3.436)	(0.124)
Oil					−0.089	−0.223**	−2.142	0.017
					(0.097)	(0.133)	(3.813)	(0.125)
Constant	−1.431***	−0.863	−66.206***	2.687***	−0.995	0.435	−40.459*	1.207
	(0.58)	(0.827)	(24.942)	(0.855)	(0.798)	(1.081)	(26.87)	(0.979)
No. of observations	56	56	58	64	49	50	51	56
R^2	0.32	0.15	0.39	0.41	0.41	0.32	0.55	0.47
F-statistic	8.0***	2.95***	11.66***	13.86***	3.41***	2.36***	6.32***	5.2***
Critical capital–labor ratio	9.1	6.8	9.0	7.8	9.3	4.2	9.2	7.5

Standard errors in parantheses; *** significant at 5% level, ** significant at 10% level, * significant at 15% level.

Our regressions help us identify the critical level (which also are estimates with standard errors) of the capital-labor ratio, for each of the measures of trade restrictions, at which, the relationship between trade restrictions and inequality changes sign. In Tables 1 and 2, we also provide this turning point or the critical capital-labor ratio. Except for the quota regressions, these numbers are fairly close to the median (9.4) and the mean (9.8) capital–labor ratios. This is specially true for the tariff and import duty regressions. In Table 3, using the tariff regression with controls presented in Table 1, we categorize the countries in our sample into those that exhibit a negative relationship between protection and inequality (those with a low capital–labor ratio) and those that exhibit a positive relationship (those with a high capital–labor ratio). The critical (turning point) capital–labor ratio in this case is roughly 8.5 which is slightly lower than the capital–labor ratio for Korea.

Table 3
Countries (tariff-gini relationship)

Negative relationship	Positive relationship
Bangladesh	Korea, South
China	Jamaica
Rwanda	Iran, Islamic Republic of
Pakistan	Venezuela
India	Portugal
Indonesia	Greece
Sri Lanka	Trinidad and Tobago
Ghana	Algeria
Nigeria	Spain
Philippines	Singapore
Thailand	Ireland
Guatemala	United Kingdom
Cameroon	New Zealand
Uganda	Malaysia
Colombia	Canada
Turkey	Italy
Chile	Belgium
Peru	Netherlands
Costa Rica	France
Tunisia	United States
Mauritius	Sweden
Jordan	Denmark
	Japan
	Finland
	Luxembourg
	Norway
	Germany
	Morocco
	Mexico

Note: The above partitioning is based on the tariff regression with controls in Table 1.

A partial derivative of trade restrictions with respect to the capital–labor ratio in the regressions with the Gini-coefficient yields

$$\frac{\partial \mathrm{TR}_i}{\partial (K/L)_i} = \alpha_3 + \alpha_2(\mathrm{INEQ}_i) \tag{7}$$

Our regression results show that $\alpha_3 < 0$ and $\alpha_2 > 0$ and their estimates are statistically significant except for the case of the quota-coverage ratio. Plugging in the values of INEQ_i into the expression for the above partial derivative, we find a negative sign overall barring very few exceptions. For example, in Table 2, the partial derivative is always negative except for Guatemala. These results are in line with the findings of Magee et al. (1989). Tariffs are a dependable and important source of revenues in developing countries (countries with a low capital–labor ratio). Moreover, developing countries have used infant-industry reasoning to justify protecting domestic industries.

We now look at the coefficients of our control variables in Tables 1 and 2. Our controls are an inverse index of democracy (the Gastil index of political rights), schooling, and regional effects using regional dummies. The inclusion of democracy is motivated by several factors. First, if we believe the evidence that openness stimulates economic growth, dictatorships which are more concerned with the size of the pie rather than its distribution, are more likely to be open. Second, since unemployment is a major issue in most elections, democracies are also more likely to provide import protection to inefficient domestic firms and to public sector firms that may not survive foreign competition. Furthermore, Fernandez and Rodrik (1991) show that in the presence of individual-specific uncertainty regarding the costs of moving to the export sector, trade reforms that are beneficial to the majority ex-post may require a dictator to implement them in the first place. Third, Rodrik (1997) has argued that rising labor demand elasticities, brought about by more open trade, may hurt workers (the majority of the population) by shifting the wage or employment incidence of non-wage labor costs towards labor and away from employers, by triggering more volatile responses of wages and employment to labor demand shocks and by shifting bargaining power over rent distribution in firms away from labor and towards capital. This may generate some demand for protection, to which democracies may be more responsive. In Table 2, we find a weak, positive link between protection and democracy in the case of quotas.[19] This is shown by the negative sign of the

[19] Additionally, we also use another control variable which is the interaction of democracy and capital–labor ratio for the reason that redistributive labor-oriented trade policies can be anti-trade or pro-trade depending on the capital abundance of the economy and democracies might be more responsive to demands for such redistribution. This variable turns out to be statistically insignificant.

coefficient of Gastil's (inverse) index of political rights.[20] Somewhat stronger results of this kind are obtained when we use alternative measures of the capital–labor ratios (discussed in detail in the next section).

In Table 2, we also find that schooling has a negative and significant effect on trade restrictions. A possible reason for this is that schooling is higher and trade restrictions lower in developed countries and that schooling is simply an index of development. Another plausible reason, however, for our observed sign could be that a better educated public is better informed about government policies and can better figure out the dead-weight costs of distortionary government policies favoring special interest groups. Finally, the inclusion of regional dummies does not affect our conclusions. East Asian economies in general seem to have had lower protection (Table 1). Apart from that, we fail to find any significant evidence whether a particular region or group of countries have a tendency to be more open or more protectionist.

5.2. Robustness checks

5.2.1. The Hausman test for endogeneity

Finally, because of the possible endogeneity of the Gini coefficient, the capital–labor ratio and the interaction between the two (as explained in the previous section), we perform the Hausman test for contemporaneous correlation between the error and the three regressors, that are suspected to be endogenous. As mentioned in the previous section, this test can be used to test any potential failure of the orthogonality assumption so long as instrumental variables are available. We use a simple method of implementing this test as suggested by Hausman (1978) and extended in Smith and Blundell (1986), the details of which have already been discussed in Section 3.

Each of our suspected endogenous variables is regressed on all the exogenous variables from our main regression plus other exogenous variables that, we believe, affect any of the three endogenous variables. The residuals from these auxilliary regressions are calculated and inserted as additional regressors in our main regression. F-Hausman in Table 4 gives the F-statistic for the joint significance of these residuals. All these residuals turn out to be jointly insignificant except in the case of the tariff. The results of the Hausman test suggest that there is no loss in consistency from the use of the OLS estimates when our dependent variable is the quota, import duty or $(X+M)/\text{GDP}$. In addition, OLS estimates are efficient. However, in the case of the tariff, the consistent coefficient estimates are only the ones obtained from the Hausman regression (identical to instrumental variable estimates) presented in Table 4. Besides, the required standard error correction to produce the IV standard errors has also been done. After taking into account the endogeneity of the capital–labor ratio, inequality and

[20] It needs to be noted that this index increases with the extent of dictatorship and decreases with the degree of democracy.

Table 4
Hausman regressions

	Tariff	Quota	Import duty	$(X+M)$/GDP
Gini	−0.171***	0.0003	−3.198***	0.043
	(0.084)	(0.055)	(1.407)	(0.05)
Gini*capital–labor ratio	0.017***	0.001	0.368***	−0.006
	(0.008)	(0.005)	(0.136)	(0.005)
Capital–labor ratio	−0.722***	−0.109	−18.15***	0.365**
	(0.326)	(0.238)	(6.148)	(0.217)
Schooling	0.007	−0.014	0.816	−0.019
	(0.047)	(0.041)	(0.938)	(0.034)
Political rights (Gastil)	0.012	0.013	1.107	0.016
	(0.049)	(0.039)	(1.013)	(0.035)
Sub-saharan Africa	−0.192	−0.43*	−5.907	−0.244
	(0.345)	(0.278)	(7.065)	(0.261)
East Asia	−0.046	−0.283**	−8.041**	0.098
	(0.215)	(0.166)	(4.214)	(0.155)
Oil	0.018	−0.03	0.02	−0.021
	(0.17)	(0.134)	(3.45)	(0.128)
Residual (Gini)	0.039	0.012	0.408	0.024
	(0.032)	(0.024)	(0.596)	(0.021)
Residual (capital–labor)	−0.013	0.062	1.153	−0.004
	(0.077)	(0.059)	(1.565)	(0.056)
Residual (capital–labor*Gini)	−0.004***	−0.001	−0.073**	−0.001
	(0.002)	(0.002)	(0.043)	(0.001)
Constant	7.25***	1.031	165.195***	−2.425
	(3.22)	(2.273)	(60.05)	(2.121)
No. of observations	37	37	37	41
R^2	0.66	0.47	0.65	0.53
F-statistic	4.52***	2.03**	4.26***	3.01***
Joint signifcance of residuals (F-Hausman)	4.34***	0.62	1.76	0.6

For tariffs, the Hausman test suggests that there is an endogeneity problem. Therefore, for this case we have modified the standard error by a factor of 1.94 to obtain consistent IV standard error estimates. After correction our main coefficient estimates remain highly significant.
The residuals are obtained from auxiliary regressions where each of the suspected endogenous variables are regressed on all exogenous variables. The additional exogenous variables are savings rate, population growth, land gini, civil liverty and $M2$/GDP
Standard errors in parantheses; *** significant at 5% level, ** significant at 10% level, * significant at 15% level.

the product of the two in the case of the tariff, we have the results virtually unchanged and still statistically significant. The critical capital–labor ratio is around 10, again fairly close to the median and the mean.

5.2.2. Dictatorship vs. democracy

We have argued earlier in this paper that majoritarian concerns are important in both democracies and dictatorships. Nevertheless, these concerns may be relatively more important in democracies. There are two possible interpretations here: (a) the

median-voter model fits the data better for democracies; and (b) the predicted relationship between trade policy and inequality is stronger (larger in magnitude) for democracies.

We investigate (a) by generating residuals from our main regressions and then regressing the absolute values and alternatively, squares of these residuals on the democracy/dictatorship (political rights) variable. In most cases, we do not get any statistically significant results, except in the case of the absolute values and squares of residuals obtained from the import-duty regression without controls (the quota residuals with the Nehru–Dhareshwar data also give us significant results, but those quota residuals are based on results which do not in the first place validate the median-voter predictions):

$$|\text{Import duty residual}| = \underset{(1.24)}{0.94} + \underset{(0.35)}{1.26}^{***}\ (\text{political rights}) \quad R^2 = 0.22$$

$$[\text{Import duty residual}]^2 = \underset{(28.45)}{1.43} + \underset{(7.99)}{15.22}^{**}\ (\text{political rights}) \quad R^2 = 0.07$$

(Note: The standard errors are shown in parentheses and asterisks represent significance). As the political rights variable is increasing in the extent of dictatorship, the above regression results at least provide some weak evidence that the median voter prediction works better in democracies than in dictatorships. We, however, find much stronger evidence when we generate predicted values of protection using our coefficient estimates and then find their correlation with the actual values separately for the dictatorship sample (countries with values of the political rights variable above 3) and the democracy sample (the rest). Using the regressions without controls, the correlation coefficients for the dictatorship sample are 0.42, 0.42, 0.3 and 0.3 for tariff, quota, import duty and $(X+M)/\text{GDP}$, respectively, while for democracies they are 0.71, 0.5, 0.8 and 0.71. The comparisons are very similar with controls.

Finally, we run regressions with additional interaction terms (Gini*pol rights and Gini*(K/L)*pol rights) to investigate the hypothesis (b) that the demand for prolabor (median-voter related) redistribution through trade policies is stronger in democracies than in dictatorships, for which we find support only with $(X+M)/\text{GDP}$ as the dependent variable. For the regression without controls, the cross-partial derivative is $\partial^2[(X+M)/\text{GDP}]/\partial\text{dictatorship}\,\partial\text{gini} = -0.013 + 0.001(K/L)$ so that democracies reinforce the positive (negative) relationship between inequality and openness in capital scarce (abundant) countries, predicted by the median voter model.[21]

5.2.3. Alternative data on capital–labor ratios

We do some robustness checks by using other data on capital per worker. Both with the Summers–Heston as well as the Easterly–Levine K/L data, it can be seen

[21]This result for $(X+M)/\text{GDP}$ is robust to the inclusion of controls. Detailed regression results can be found at www.ualberta.ca/economics/dm/dm.htm.

Table 5
Gini coefficient-regression with controls

	Summers–Heston capital–labor ratio				Easterly–Levine capital–labor ratio			
	Tariff	Quota	Import duty	$(X+M)$/GDP	Tariff	Quota	Import duty	$(X+M)$/GDP
Gini	−0.144***	0.021	−6.268***	0.126*	−0.152*	−0.035	−6.461***	0.352
	(0.059)	(0.075)	(1.877)	(0.082)	(0.095)	(0.105)	(2.464)	(1.788)
Gini*capital–labor ratio	0.014***	−0.003	0.666***	−0.015**	0.015*	0.003	0.675***	−0.047
	(0.006)	(0.008)	(0.199)	(0.009)	(0.01)	(0.011)	(0.258)	(0.186)
Capital–labor ratio	−0.936***	−0.067	−37.324***	0.714**	−0.963***	−0.289	−37.519***	4.639
	(0.256)	(0.33)	(8.444)	(0.364)	(0.389)	(0.446)	(10.515)	(7.685)
Schooling	0.034*	−0.017	1.29**	−0.004	0.036*	−0.014	1.249*	−0.93**
	(0.021)	(0.03)	(0.749)	(0.032)	(0.022)	(0.032)	(0.772)	(0.555)
Political rights (Gastil)	−0.005	0.098***	0.287	0.003	−0.001	0.101***	0.395*	0.33
	(0.032)	(0.043)	(1.182)	(0.048)	(0.034)	(0.044)	(1.224)	(0.83)
Sub-saharan Africa	−0.346**	−0.231	−5.594	0.341	−0.334**	−0.198	−4.9	2.727
	(0.187)	(0.246)	(6.2)	(0.289)	(0.199)	(0.256)	(6.377)	(4.932)
East Asia	−0.254**	−0.375***	−6.682	−0.01	−0.245*	−0.343**	−6.003	1.697
	(0.142)	(0.187)	(4.964)	(0.22)	(0.154)	(0.196)	(5.242)	(3.797)
Oil	0.208*	0.175	10.033***	−0.025	0.223*	0.147	10.527***	−2.836
	(0.134)	(0.177)	(4.453)	(0.205)	(0.148)	(0.193)	(4.742)	(3.594)
Constant	9.253***	1.236	354.941***	−5.909**	9.595***	3.33	361.65***	−33.71
	(2.414)	(3.082)	(78.978)	(3.462)	(3.723)	(4.202)	(100.479)	(73.884)
No. of observations	31	31	31	35	28	28	28	32
R^2	0.7	0.63	0.73	0.47	0.7	0.61	0.75	0.15
F-statistic	6.29***	4.73***	7.57***	2.89***	5.67***	3.68***	7.02***	0.5
Critical capital–labor ratio	10.3	7.0	9.4	8.4	10.1	11.7	9.6	7.5

Standard errors in parantheses; *** significant at 5% level, ** significant at 10% level, * significant at 15% level.

Table 6
Third quintile-regression with controls

	Summers–Heston capital–labor ratio				Easterly–Levine capital–labor ratio			
	Tariff	Quota	Import duty	$(X+M)$/GDP	Tariff	Quota	Import duty	$(X+M)$/GDP
$Q3$	0.321***	0.191*	10.698***	−0.167	0.293***	0.21*	9.841***	−1.649
	(0.087)	(0.125)	(3.024)	(0.135)	(0.125)	(0.13)	(3.416)	(2.224)
$Q3$*capital–labor ratio	−0.034***	−0.025**	−1.226***	0.021	−0.032***	−0.029***	−1.14***	0.231
	(0.01)	(0.014)	(0.33)	(0.015)	(0.013)	(0.014)	(0.364)	(0.239)
Capital–labor ratio	0.272***	0.365**	12.756***	−0.175	0.345**	0.445***	11.497***	−1.302
	(0.129)	(0.185)	(4.414)	(0.196)	(0.19)	(0.2)	(5.152)	(3.409)
Schooling	0.005	−0.039	−0.436	0.021	0.002	−0.025	−0.401	−0.962**
	(0.022)	(0.031)	(0.736)	(0.032)	(0.03)	(0.031)	(0.81)	(0.514)
Political rights (Gastil)	−0.051**	−0.053	−2.562***	0.006	−0.031	−0.055*	−2.266***	0.45
	(0.026)	(0.038)	(0.957)	(0.04)	(0.034)	(0.036)	(1.026)	(0.521)
Sub-saharan Africa	−0.208**	−0.029	3.978	0.143	−0.015	0.079	2.234	0.19
	(0.117)	(0.166)	(3.995)	(0.179)	(0.178)	(0.18)	(5.004)	(3.066)
East Asia	−0.124	−0.002	0.853	−0.025	−0.007	0.04	−0.103	1.205
	(0.132)	(0.188)	(4.594)	(0.204)	(0.178)	(0.186)	(5.04)	(3.143)
Oil	−0.036	−0.455**	−2.222	−0.104	−0.028	−0.465**	−2.017	−1.242
	(0.177)	(0.254)	(6.126)	(0.276)	(0.234)	(0.246)	(6.406)	(4.213)
Constant	−2.215**	−2.059	−88.009***	1.476	−2.874*	−2.731*	−75.765*	10.326
	(1.201)	(1.722)	(41.302)	(1.838)	(1.813)	(1.892)	(48.77)	(32.057)
No. of observations	33	34	35	38	30	31	31	36
R^2	0.65	0.43	0.66	0.48	0.44	0.44	0.67	0.15
F-statistic	5.59***	2.36***	6.22***	3.39***	2.08***	2.16***	5.67***	0.62
Critical capital–labor ratio	9.4	7.6	8.7	8.0	9.2	7.2	8.6	7.1

Standard errors in parantheses; ***significant at 5% level, **significant at 10% level, *significant at 15% level.

from Tables 5 and 6 that our main results remain qualitatively unchanged. Inequality, the capital–labor ratio and the product of the two are very significant and have the right signs. This is true even for quotas when the third quintile's share is used as an inverse measure of inequality (Table 6). *So, there is, at least, some weak evidence for the median voter prediction working in the case of quotas*. The critical capital–labor ratios presented in Tables 5 and 6 are in most cases quite close to the mean and median values. Again, there is some evidence that more democracy leads to more protection (Table 6). The results with schooling are somewhat mixed.

5.2.4. Regressions using changes in protection and changes in inequality

We also perform regressions of changes in import duty and alternatively, changes in $(X+M)/\text{GDP}$ (the other protection measures being purely cross-sectional, i.e. available only at one point in time) on changes in inequality and an interaction term of change in inequality interacted with the capital–labor ratio. Using Δ to denote changes, Eq. (6) from Section 3 can be written as $\Delta\text{TR}_i/\Delta\text{INEQ}_i = \alpha_1 + \alpha_2(K/L)_i$ which in turn gives us our following new estimating equation:

$$\Delta\text{TR}_i = \alpha_1\,\Delta\text{INEQ}_i + \alpha_2(K/L)_i\,\Delta\text{INEQ}_i + e_i \tag{8}$$

The changes in these variables are for the 1980 to 1990 period for each country, while the data on K/L as before are the averages for the 1980s. The result for import duty seems to support and strengthen our earlier results. For $(X+M)/\text{GDP}$ we failed to find any significant relationship, perhaps because the change in GDP dominates and is itself driven by extraneous factors. The following are the regression results with the change in import duty as the dependent variable, with and without the constant term, respectively (standard errors are shown in parentheses):

$$\Delta(\text{import duty}) = \underset{(0.74)}{0.55} - \underset{(0.89)}{2.21}^{****}\,\Delta(\text{gini}) + \underset{(0.096)}{0.224}^{***}(K/L)\,\Delta(\text{gini})$$

$$R^2 = 0.18,\ N = 33$$

$$\Delta(\text{import duty}) = -\underset{(0.88)}{2.31}^{****}\,\Delta(\text{gini}) + \underset{(0.094)}{0.233}^{***}(K/L)\,\Delta(\text{gini})$$

$$R^2 = 0.18,\ N = 33$$

The signs of the coefficients of $\Delta(\text{gini})$ and $(K/L)\,\Delta(\text{gini})$ are negative and positive respectively, exactly as predicted. The critical value of the K/L ratio is 9.8, again very close to the median.

6. Conclusion

The prediction of the median voter approach (within a Heckscher–Ohlin framework) is that trade policies will be biased towards trade in labor rich countries, and biased against trade in capital rich economies. However, trade policies, as we know, are always and everywhere biased against trade. This paper gives a second chance to the median-voter approach by focusing on cross-country variations in (rather than the orientations of) trade policies. The data show that an increase in inequality increases import protection in capital-abundant countries, but reduces trade barriers in capital-scarce economies. This is consistent with the predictions of the median-voter approach within a two-factor, two-sector Heckscher–Ohlin model. There is some evidence, that this relationship may hold better in democracies than in dictatorships.

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Appendix A

Here we tie our empirical results to the theoretical predictions in the context of the inequality measures used. Let the share of the median quintile in the capital stock be λ_K and its share in total income be λ_I. Since in the model, tariff revenue is distributed in proportion to a person's share in factor income, in the context of our two-factor (capital–labor) framework, we have $\lambda_I = [0.2wL + r\lambda_K K]/[wL + rK] = [0.2w + r\lambda_K k]/[w + rk]$ where $k = K/L$ is the economy's ratio of capital to labor endowment. $\partial\lambda_I/\partial\lambda_K = \frac{rk}{w+rk} > 0$, $\partial\lambda_I/\partial p = (0.2 - \lambda_K)k[rw'(p) - wr'(p)]/[w + rk]^2 > 0$ for a capital-abundant country and < 0 for a labor abundant country (as in any country with asset inequality, $\lambda_K < 0.2$). Thus, we can write $\lambda_I = \psi(\lambda_K, t, k)$ where t is the tariff and $\psi_1 > 0$, and $\psi_2 > 0$ for a capital-abundant country and < 0 for a capital-scarce economy. Holding λ_I constant and allowing λ_K and t to

vary, we have $d\lambda_I = 0 \Rightarrow \psi_1 \partial\lambda_K/\partial t + \psi_2 = 0 \Rightarrow \partial\lambda_K/\partial t = -\psi_2/\psi_1 < 0$ for a capital-abundant country and > 0 for a capital-scarce country. Thus, we can write $\lambda_K = \phi(\lambda_I, t, k)$ where $\phi_1 > 0$, and $\phi_2 < 0$ for a capital-abundant country and > 0 for a capital-scarce one. The equilibrium tariff can be written as $t = f(k, \lambda_K)$ where $f_2 < 0$ for a capital-abundant country and > 0 for a capital-scarce economy. Therefore, we have $t = f(k, \phi(\lambda_I, t, k)) \Rightarrow dt/d\lambda_I = f_2\phi_1/[1 - f_2\phi_2]$. We find from our regressions that $f_2\phi_1/[1 - f_2\phi_2] < 0$ for a capital-abundant country and > 0 for a capital-scarce one. We know that with the two-factor (capital–labor) Stolper–Samuelson effect at work, we have support for the median-voter predictions only if $f_2 < 0$ for a capital-abundant country and > 0 for a capital-scarce one. The question here is whether, with the Stolper–Samuelson effect at work, the signs of the estimated derivative $dt/d\lambda_I$ from our regressions support the median-voter predictions. In a capital−abundant country $f_2\phi_1/[1 - f_2\phi_2] < 0 \Rightarrow$ either $\{\underset{(-)(+)}{f_2\phi_1} < 0, \underset{(-)(-)}{1 - f_2\phi_2} > 0\}$ or $\{\underset{(+)(+)}{f_2\phi_1} > 0, \underset{(+)(-)}{1 - f_2\phi_2} < 0\}$ (Note that we have already established $\underset{(-)(-)}{\phi_2}$'s negative sign for a capital-abundant country and ϕ_1's positive sign in general). Only $\{\underset{(-)(+)}{f_2\phi_1} < 0, \underset{(-)(-)}{1 - f_2\phi_2} > 0\}$ consists of mutually consistent inequalities. Similarly when we find that $f_2\phi_1/[1 - f_2\phi_2] > 0$, for a capital-scarce economy we have $\{\underset{(+)(+)}{f_2\phi_1} > 0, \underset{(+)(+)}{1 - f_2\phi_2} > 0\}$ or $\{\underset{(-)(+)}{f_2\phi_1} < 0, \underset{(-)(+)}{1 - f_2\phi_2} < 0\}$, of which only the first set consists of mutually consistent inequalities. Thus, our derivatives from our regressions imply that $f_2 < 0$ (> 0) for a capital-abundant (capital-scarce) country. Thus, finding a positive (negative) relationship between inequality and protection in a capital-abundant (scarce) economy provides support for the median-voter predictions.

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