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# Does inequality impact tax collection? Evidence from ACI (ASEAN-China-India) economies

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### *Abstract*

This study examines the short-run and the long-run relationships between inequalities – measured by the (income) Gini coefficient – and taxes, using a panel of ten selected Asian countries from 1993 to 2015. After testing for the applicability of several econometric models of the panel Autoregressive Distributed Lag (ARDL) methodology, we choose the Pooled Mean Group (PMG) estimator to find that an increase in income Gini increases the tax-GDP ratio for the ACI economies in the long run. However, we also note that the income Gini has no (statistically significant) effect on taxes in the short run. The chain of causality is found to run from income inequalities to taxes and not from taxes to inequalities. This study confirms the prediction of the median voter hypothesis on the consequences of income distribution: greater inequality is associated with a larger tax-GDP ratio because of the greater redistribution that is sought by the median voter when income distribution is less equal.

**Keywords**: Inequality; Taxes; Pooled Mean Group (PMG) Autoregressive Distributed Lag (ARDL)

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## 1. INTRODUCTION

Over the period 1990-2015 remarkable economic achievements were recorded in Asia – despite the Asian Financial Crisis and the Global Financial Crisis rocking their regional economies – as the region grew at 6% per year (see Jain-Chandra et al., 2016). The poverty rate declined from 55% in 1990 to about 20% in 2015 (Jain-Chandra et al., 2016). Against this backdrop of economic successes, the Asian economies also witnessed rising income inequalities: since 1990 growth in the average Gini coefficient has been higher in Asia than for the rest of the world (see Zhuang, Kanbur & Maligalic, 2014, pp. 32 and 34, Figure 2.8). Further increases in income inequalities over this period impacted Asian inequalities as the population-weighted Gini for Asia rose from 39 to 46 (see Zhuang, Kanbur & Maligalic, 2014, p. 36). The populous countries of ASEAN, along with China and India, have experienced continuing increases in income inequalities in the region, which motivates us to examine the precise impact of inequalities on taxes for the populous nations of ASEAN, China and India (collectively called the ACI economies).

With the unprecedented GDP growth, ACI economies experienced significant technological changes, increases in labour force participation by low-skilled workers, declining top marginal income tax rates, widening inter-regional inequality within their economies and pressures from globalisation and liberalisation of regional factors and product markets. These changes are held responsible for growing inequality (see IMF, 2014). The impacts of inequality on other economic outcomes have been extensively studied in the extant literature.[1] In this article, we seek to establish whether income Gini can drive fiscal (tax) outcomes. Our primary motivation behind this work is predicated upon the possibility that inequality can significantly influence the political outcomes and, thereby, fiscal outcomes in a society. It is an accepted tenet of public finance that governments exist for the provision of public goods in addition to fighting adverse consequences of missing markets, imperfect information and externalities (see Grossman, 1988). Governments also offer growth stimuli by laying down public investment that, in turn, promotes productivity of private investment (Khan & Kumar, 1997). Thus, effective governance calls forth adequate resources at the command of governments, whether in developing or developed countries, 'to satisfy not only the short-term needs of its population but also its long-term developmental goals' (Tran-Nam & Le, 2022, p. 194). One of the main development goals is to promote equitable distribution to fight poverty (see Goda, 2017; Galor & Moav, 2004; Rodrik, 1999 among others).

The increasing relevance of governments in modern societies is inexorably linked with Wagner's Law which posits that government spending, and hence tax revenue, is

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[1] Though income inequalities might provide incentives for investment and, thereby, trigger economic growth (Barro, 2000; Forbes, 2000), income inequalities are also held responsible for adversely impacting on macroeconomic stability and sustainable growth (Ostry, Berg & Tsangarides, 2014). Cingano (2014) found that higher inequality fosters aggregate savings which permits capital accumulation because the rich have a lower propensity to consume. Galor and Moav (2004) showed that rising inequality can compromise the health status of the poor and formation of human capital and thereby undermine growth. Alesina and Perotti (1995) and Perotti (1996) argued that political and economic instabilities – caused by rising inequality – reduce investment and, hence, reduce growth. The work of Mah-Hui and Khor (2011) and Goda (2017) showed that rising inequalities can trigger financial shocks. Rodrik (1999) argued that inequalities reduce social harmony that is necessary to maintain resilience by absorbing economic and financial shocks.

endogenous and positively responds to the rising per capita income of a country (Peacock & Scott, 2000). In other words, economic growth paved the way for 'cultural and economic progress' such that the public demands larger state activities in lieu of private economic activities (Peacock & Wiseman, 1961, p. 16). Our main goal in this article is to investigate the effect of inequalities on taxes.

Despite a relative scarcity of studies on the impacts of inequalities on taxes, in an interesting recent work, Islam et al. (2018) utilise the panel models to examine the effects of inequalities on income tax-GDP ratios for 21 OECD countries over the period 1870-2011. Since taxes have been used by policy-makers to reduce inequalities (Islam et al., 2018), the reverse causality running from taxes to inequalities can create estimation problems for any model examining the impact of inequality on taxes. This is an important element missing from the work of Islam et al. (2018). In this article, we entertain the possibility of mutual causality, or interdependency, between inequalities and taxes. We will then establish that there is no evidence of reverse causality running from taxes to inequalities for the countries of our choice. Hence, our work will establish unequivocally whether inequalities impact taxes for the ACI economies. It is also imperative to note that increases in within-country income inequalities in the ACI region, during 1990-2015, were the largest in the global economy. The plan of the remainder of the article is as follows: in section 2, we review the relevant literature. Section 3 outlines the analytical foundation, methodology, and data sources. In section 4, we discuss the findings. Finally, in section 5, we conclude.

## 2. LITERATURE REVIEW

As governments seek to control inequality through fiscal measures, several researchers have studied the determinants of endogenous tax policies and taxes (e.g., Hettich & Winer, 1988; Besley & Case, 1995; Milanovic, 2000; Harms & Zink, 2003; Aidt & Jensen, 2009; Corneo & Neher, 2015). Although only a few studies have explored the impact of inequality on taxes, the findings are mixed (Islam et al., 2018): for instance, in an analysis of 50 countries for the period 2007-2011, Aizenman and Jinjarak (2012) note that tax revenues fall as a percentage of GDP as inequality rises. However, analysing data from 75 countries for the year 2004, Adam, Kammas and Lapatinas (2015) show that as inequality increases, (capital) taxes as a percentage of GDP increase, while the share of labour taxes declines. At the regional level, Boustan et al. (2013) show that inequality increases taxes in the US – at the municipality and school district level from 1970 to 2000. By contrast, using a laboratory setting, Agranov and Palfrey (2015) argue that inequality increases tax rates.

Analysing the impact of income inequality on the tax capacities of 96 countries using the tax stochastic frontier model, Pessino and Fenochietto (2010) confirm that a higher Gini coefficient lowers the tax capacities of governments and thus adversely impacts tax efforts. In a subsequent analysis, Fenochietto and Pessino (2013) analysed 113 countries and concluded that (among other factors) income distribution (Gini coefficient) negatively impacted tax revenues as a percentage of GDP. Additionally, the study found that European economies with strong income distribution policies operated near their tax capacity.

According to Bird, Martinez-Vazquez and Torgler (2014), if income inequality results from the unfair distribution of tax burdens, the consequence would be lower levels of trust in public institutions, eventually lowering tax efforts.[2] Similarly, treating income inequality as a factor explaining tax inefficiency in a Stochastic Tax Frontier Model, Tran-Nam and Le (2022) find that the taxpayers' perception of a higher inequity increases the tax level non-compliance via several tax-evading and avoidance measures.

In an analysis of the issue in 21 OECD countries, contrary to the predictions of the median voter argument,[3] Islam et al. (2018) noted that inequality depresses income tax ratios and suggest that the effect is more significant for more democratic countries.[4] However, in a subsequent analysis of OECD countries using 'Social Inequality cumulation', Kuhn (2019) argues that voters, perceiving a high level of wage inequality, tend to be more supportive of redistributive policies and progressive taxation.

## 3. MOTIVATION, METHODOLOGY AND DATA

In what follows in section 3.1 below, we provide the rational foundation of our work and the main motivation behind this study.

### 3.1 Political transaction costs: inequality *vis-à-vis* taxes

We make a departure from the analysis in previous studies by exploiting the concept of political transaction costs[5] to understand and uncover the precise relationships between taxes and inequalities. For political markets – being characterised by incomplete political rights, imperfect enforcement agreements, bounded rationality, and imperfect information (as highlighted by the New Institutional Economics: see North, 1990a, 1990b; Nye, 1997; Pierson, 2000; Moe, 2005, among others) – institutions, conventions, and rules of the game become crucial determinants of the political inputs and political outputs due to pervasive transaction costs (see Pierson, 2000). It is well-received in the political transaction cost literature that determinants of political outcomes are often 'opaque' and 'unclear' with limited observability and measurability (Pierson, 2000). As reflected in an earlier work of Dahlman (1979), the relevant political transaction costs of tax policies can assume three forms: first, search and information costs of suitable tax policies. Secondly, bargaining and negotiation costs among political actors. Finally, the enforcement, monitoring and policing costs. If inequalities impinge on the determinants of transaction costs, then inequalities can influence tax outcomes. The primary motivation of this study is to empirically assess the precise relationship between inequalities and tax outcomes.

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[2] Tax effort is defined as the ratio of actual taxes and the potential taxes (Tran-Nam & Le, 2022).

[3] The median voter model implies that the political outcomes in a democracy reflect the median voter's preference (Congleton, 2004). The theory predicts that, under political pluralism, political parties compete for the majority of the voters by focusing their attention on the outcome most preferable by an electorate with a median income. Thus, an expected result is that as income inequality expands, there will be an increase in taxes to serve the distributive interest of the median voter.

[4] These authors claim that a long term decline in market inequality, especially between 1915 and 1980, resulted in historical growth in tax rates; however, since the 1980s, rising inequality has adversely impacted tax revenues, exerting pressure on national budgets and debts.

[5] The role of transaction costs in the political process has been amply demonstrated in the work of North (1990a) and Dixit (1996, 2003): the transaction cost theory in politics is predicated upon (i) costly information, (ii) subjective models of decision-making by political actors, and (iii) imperfect enforcement of agreements (see North, 1990a; p. 355). Such costs arise for transactions between politicians and citizens (Dixit, 1996, 1998). Political transaction costs are also rampant in interactions between politicians (see Weingast & Marshall, 1988; Epstein & O'Halloran, 1999; Spiller & Tommasi, 2007).

## 3.2 Methodology

Our methodology is based on a positivist research framework and a quantitative method utilising secondary data. More specifically, we postulate a simple model that taxes are determined by inequalities as applied by Islam et al. (2018) and others:

$$\ln TAX_{it} = \beta_0 + \beta_1 \ln TAX_{i,t-1} + \beta_2 \ln GINI_{it} + \gamma X_{it}{}' + a_i + \varepsilon_{it}, \quad \textbf{(1a)}$$

where $TAX_{it}$ is the share of tax revenues in GDP for country $i$ in year $t$, $GINI_{it}$ is the measure of income inequality for country $i$ in year $t$, $X_{it}$ is a vector of control variables – namely urbanisation ($\ln URBAN_{it}$), per capita GDP ($\ln PCGDP_{it}$) and trade openness ($\ln TRADE_{it}$) – ln is the natural logarithmic transformation of the variables (not the decimal values). Note that $a_i$ denotes the country fixed effect, and $\varepsilon_{it}$ is the error term.

The panel ARDL model, which we will apply, is specified in section 3.3.2. In section 3.3.1, we summarise the panel unit root tests to justify the rationale for using the panel ARDL methodology. Islam et al. (2018), instead of applying the panel ARDL methodology, used panel data analysis, which fails to adequately handle the non-stationarity of variables (like taxes and inequalities) over a long haul with the possibility of spurious statistical significance (see Brückner & Ciccone, 2011; Ciccone, 2011, 2013). The panel ARDL method is robust to handle both autocorrelation, and non-stationarity (see Alsamara et al., 2017; Gangopadhyay & Nilakantan, 2018), and can simultaneously handle both stationary and nonstationary variables, thus bypassing the critique of Ciccone (2013).

## 3.3 Variables and data

Our panel sample has ten (10) countries, comprising eight (8) populous nations (Myanmar, Cambodia, Indonesia, Laos, Malaysia, the Philippines, Thailand and Vietnam) of the ten countries from the ASEAN grouping and China and India, for the 23 years of the period 1993-2015. We ignored Singapore and Brunei for their small populations.

The ACI economies have a total population of about 3 billion and a combined GDP of USD 4.8 trillion. The average tax to GDP ratio among our sample countries is 12.12%, and the mean of our measure of the distribution of income, the Gini coefficient, is 34.80. Countries analysed in the study have average urbanisation of 36.28%, and trade openness (ratio of total trade to GDP) for the samples stood at 86.64%.

**Table 1: Summary Statistics for the ACI Economies from 1993 to 2015**

| Variable | Mean | Standard Deviation | Minimum | Maximum |
|---|---|---|---|---|
| Tax to GDP (TAX) | 12.12 | 4.90 | 1.96 | 22.46 |
| Gini (GINI) | 34.80 | 6.32 | 25.65 | 53.26 |
| Urbanisation (URBAN) | 36.28 | 13.71 | 16.49 | 74.21 |
| Trade Openness (TRADE) | 86.64 | 48.31 | 19.31 | 220.41 |
| Per Capita GDP (lnPCGDP) | 8.95 | 5 | 5.41 | 23.04 |

Data Source: The Asian Development Bank (ADB) website; inequality dataset also uses the WIID database of UNU-WIDER to have a consistent series.

### *3.3.1 Panel unit root tests*

Prior to conducting any estimations, panel unit root tests were implemented to assess the order of integration. The Levin, Lin and Chu (2002) (LLC) test and the Im, Pesaran and Shin (2003) (IPS) test are the two most extensively employed techniques to determine the stationarity of variables in panel studies. While the LLC test results depend on pooled data, the IPS test results are based on the average of Augmented Dickey-Fuller (ADF) statistics.

The LLC panel unit root test for each variable of interest $y_i$ is based on the following equation:

$$y_{it} = \rho_i y_{i,t-1} + z_{it}'\gamma + u_{it} \qquad \textbf{(1b)}$$

$i = 1\ldots, N; t = 1, T$

where $z_{it}$ is the deterministic component and $u_{it}$ is a stationary process. One of the assumptions of the LLC test is that residuals are independently and identically distributed with zero mean and variance $\sigma_u^2$ and $\rho_i = \rho$ for all values of i. The null hypothesis of the LLC test is, $H_0: \rho = 1$, which means that all series in the panel have a unit root, whereas the alternative is $H_1: \rho < 1$, which means that all series are stationary (Bildirici, 2014).

While the LLC test allows for heterogeneity in the intercept terms, the IPS test allows for heterogeneity in both the slope and the intercept terms for the cross-section units. The IPS unit root test can be specified as:

$$y_{it} = \rho_i y_{i,t-1} + \sum_{j=1}^{p_i} \varphi_{ij}\, \Delta y_{i,t-j} + z_{it}'\gamma + \varepsilon_{it} \qquad \textbf{(2)}$$

Similarly to LLC, IPS tests the null hypothesis $H_0: \rho = 1$, which means that all series in the panel have a unit root. The alternative hypothesis of the test is that part of the series is stationary, i.e. $H_1: \rho < 1$.

Table 2 below describes the panel unit root test results for our key variables of interest.

**Table 2: Unit Root Test Results for the Variables of Interest**

| Variables | Levin, Lin & Chu | | Im, Pesaran & Shin | |
|---|---|---|---|---|
| | Level | First Difference | Level | First Difference |
| lnTAX | -1.692*** | -5.325*** | -0.7573 | -4.577*** |
| lnTRADE | -0.966 | -4.663*** | 1.512 | -4.705*** |
| lnURBAN | -4.172*** | -1.717** | -1.291* | 1.306 |
| lnPCGDP | -2.493*** | -3.571*** | -0.5527 | -2.951*** |
| lnGINI | -4.338*** | -1.959** | -2.870*** | -2.620*** |

*** indicates significance at the 1% level, ** indicates significance at the 5% level, and * indicates significance at the 10% level. Lag lengths were determined by the Schwarz Information Criterion (SIC).

The panel unit root results in Table 2 indicate that the variables of interest are stationary at either level or first difference, confirming that the panel ARDL technique is suitable for our study. Precisely, the LLC test shows that lnTAX, lnURBAN, lnPCGDP, and lnGINI are stationary at $I(0)$, while lnTRADE is stationary at $I(1)$. The IPS unit root test indicates that lnURBAN and lnGINI are stationary at level, while all other variables of interest are stationary at first difference.

### 3.3.2 *Panel ARDL models*

A major issue in our dataset is that not all our variables of interest are integrated of the same order. To overcome this problem, we have employed the panel ARDL technique, as proposed by Pesaran and Smith (1995) and Pesaran et al. (1999). The Panel ARDL model is a variety of the ARDL (p,q) model, which is estimated as below (Pesaran et al., 1999):[6]

$$y_{it} = \sum_{j=1}^{p} \lambda_{ij} y_{i,t-j} + \sum_{j=0}^{q} \delta'_{ij} x_{i,t-j} + \mu_i + \varepsilon_{it} \quad \textbf{(3)}$$

where $y_{it}$ is the dependent variable (Taxes) $x_{it}$; $(k \times 1)$ is the vector of explanatory variables for group i (Gini and control variables); $\mu_i$ represents the fixed effects; the coefficients of the lagged dependent variables, $\lambda_{ij}$, are scalars; and $\delta_{ij}$ are $k \times 1$ coefficient vectors. $\varepsilon_{i,t}$ represents the error terms, $i\ (= 1,2, \ldots, N)$ labels country i and $t = (1,2, \ldots, T)$ labels year t.

[6] Panel ARDL is a preferred option if panel cointegration models are not applicable for regressors being I(0) and I(1). Pesaran and Shin (1999) argued that the method of panel ARDL is superior regardless of whether the underlying regressors exhibit I(0), I(1) or a mixture of both and the time span (T) is over 20 years and number of panels (N) is small. It is not appropriate to use the dynamic generalised method of moments (GMM) estimators due to the nature of dataset with N=10 and T=23. Following the extensive literature on dynamic panel data, we will implement several estimators to assess the postulated relationship between taxes and inequalities, by assessing the suitability of Mean Group (MG), Pooled Mean Group (PMG) and Dynamic Two-Way Fixed Effect (DFE) estimators (see Pesaran & Smith, 1995; Pesaran et al., 1999).

The re-parametrised ARDL $(p, q, q, \ldots, q)$ error correction model is specified as:

$$\Delta y_{i,t} = \Phi_i y_{i,t-1} + \beta_i' x_{it} + \sum_{j=1}^{p-1} \lambda^*{}_{ij} \; \Delta y_{i,t-j} + \sum_{j=0}^{q-1} \delta_{ij}^* \, \Delta x_{i,t-j} + \mu_i + \varepsilon_{i,t} \qquad \textbf{(4)}$$

Note in (4):

i = 1,2,…,N, and t = 1,2,…,T ,

$\Phi_i = -(1 - \sum_{j=1}^{p} \lambda_{ij})$ ,

$\beta_i = \sum_{j=0}^{q} \delta_{ij}$,

$\lambda^*{}_{ij} = -\sum_{m=j+1}^{p} \lambda_{im}$ , $\quad j = 1,2, \ldots, p-1$,

$\delta^*{}_{ij} = \sum_{m=j+1}^{q} \delta_{im}$ , $\quad j = 1,2, \ldots, q-1$.

If we stack the time-series observations for each sample, then (4) can be written as:

$$\Delta Y_i = \Phi_i Y_{i,-1} + X_i \beta_i + \sum_{j=1}^{p-1} \lambda_{ij}^* \Delta Y_{i,-j} + \sum_{j=0}^{q-1} \Delta X_{i,-j} \delta_{ij}^* + \mu_i + \varepsilon_i \qquad \textbf{(5)}$$

$i = 1,2, \ldots, N$, where $Y_i = (y_{i1}, \ldots, y_{iT})'$ is a $T \times 1$ vector of the observations on the dependent variable of the ith group, $X_i = (x_{i1}, \ldots, x_{iT})'$ is a $T \times k$ matrix of observations on the regressors that vary both across groups and time periods, $l = (1, \ldots, 1)'$ is a $T \times 1$ vector of 1s, $Y_{i,-j}$ and $X_{i,-j}$ are j period lagged values of $y_i$ and $\Delta X_i$, and $\Delta Y_i = Y_i - Y_{i,-1}$, $\Delta X_i = X_i - X_{i,-1}$, $\Delta Y_{i,-j}$ and $\Delta X_{i,-j}$ are j period lagged values of $\Delta Y_i$ and $\Delta X_i$, and $\varepsilon_i = (\varepsilon_{i1}, \ldots, \varepsilon_{iT})'$.

### *3.3.3 Pooled Mean Group (PMG) estimator*

When analysing panel data, econometric approaches can be classified into two distinct categories, namely, the Mean Group (M.G.) estimator and the Pooled Mean Group (PMG) estimator. Proposed by Pesaran and Smith (1995), the M.G. estimator accommodates individual heterogeneity by estimating individual equations for each cross-section and averaging the parameter estimates. This might appear to be a consistent estimator, yet it is not necessarily an efficient estimator of the average of the heterogeneous parameters. Alternatively, the cross-sections can be pooled, which allows for different intercepts but requires that the slope parameters are alike, which is regarded as a highly restrictive assumption (Asafu-Adjaye et al., 2016).

The PMG estimator offers a balance between these two competing approaches. It allows short-run coefficients to vary across countries (like the M.G. estimator), while the long-run coefficients are required to be homogeneous for all cross-sections (akin to the fixed effects estimator). Some of the key advantages of the PMG estimator are: first, the PMG estimator can be engaged to analyse the variables regardless of whether the variables are I(0) or I(1) – as is the case in the present study; second, short-run causality inferences can be drawn even if the presence of cointegration is not formally detected, and finally, if variables are in logarithms, then the long-run coefficients can be interpreted as elasticities (Pesaran et al., 1999).

Consider the following ARDL (1,0,2,0,0) equation for the income Gini $Y_{it}$ for country i at time t:

$$Y_{it} = \lambda_i Y_{i,t-1} + \sum_{j=0}^{1} \delta'_{ij} X_{i,t-j} + \mu_i + \varepsilon_{it} \quad \textbf{(6)}$$

where $X_{i,t-j}$ is an $n \times k$ vector of the logarithms of the explanatory variables (lnGINI, lnTRADE, lnURBAN, lnPCGDP), $\delta_{ij}$ is a $k \times 1$ coefficient vector and $\mu_i$ accounts for country-specific effects. Equation (6) above can be rearranged into an error correction model of the following form:

$$\Delta Y_{it} = \emptyset_{1,i}\left(Y_{i,t-1} - \theta'_{1,i} X_{i,t-1}\right) + \delta^{*\prime}_{1,i}\, \Delta X_{it} + \mu_i + \varepsilon_i \quad \textbf{(7)}$$

Similarly, the remaining equations can be expressed – in a panel vector autoregressive (VAR) framework using variable names – as the following:

$$\Delta \mathrm{lnTAX}_{it} = \emptyset_{1,i}\left(\mathrm{lnTAX}_{i,t-1} - \theta'_{1,i} X_{i,t-1}\right) + \delta^{*\prime}_{1,i}\, \Delta X_{it} + \mu_i + \varepsilon_i \quad \textbf{(8.1)}$$

$$\Delta \mathrm{lnGINI}_{it} = \emptyset_{1,i}\left(\mathrm{lnGINI}_{i,t-1} - \theta'_{1,i} X_{i,t-1}\right) + \delta^{*\prime}_{1,i}\, \Delta X_{it} + \mu_i + \varepsilon_i \quad \textbf{(8.2)}$$

$$\Delta \mathrm{lnTRADE}_{it} = \emptyset_{1,i}\left(\mathrm{lnTRADE}_{i,t-1} - \theta'_{1,i} X_{i,t-1}\right) + \delta^{*\prime}_{1,i}\, \Delta X_{it} + \mu_i + \varepsilon_i \quad \textbf{(8.3)}$$

$$\Delta \mathrm{lnURBAN}_{it} = \emptyset_{1,i}\left(\mathrm{lnURBAN}_{i,t-1} - \theta'_{1,i} X_{i,t-1}\right) + \delta^{*\prime}_{1,i}\, \Delta X_{it} + \mu_i + \varepsilon_i \quad \textbf{(8.4)}$$

$$\Delta \mathrm{lnPCGDP}_{it} = \emptyset_{1,i}\left(\mathrm{lnPCGDP}_{i,t-1} - \theta'_{1,i} X_{i,t-1}\right) + \delta^{*\prime}_{1,i}\, \Delta X_{it} + \mu_i + \varepsilon_i \quad \textbf{(8.5)}$$

where, in each instance, $X_{i,t}$ is an $n \times k$ vector of the remaining explanatory variables; the short-run coefficients are denoted by $\delta^*$s and the $\theta$s denote long-run coefficients; and the $\emptyset$s represent the panel error correction terms (ECTs), which must be both negative and significant to confirm the existence of a long-run relationship between the dependent and explanatory variables (Pesaran et al., 1999).

**Table 3: The Pooled Mean Group Results for the ACI Economies**

| Dependent Variable | Model 1 lnTAX (Eq: 8.1) | Model 2 lnGINI (Eq: 8.2) | Model 3 lnTRADE (Eq: 8.3) | Model 4 lnURBAN (Eq: 8.4) | Model 5 lnPCGDP (Eq: 8.5) |
|---|---|---|---|---|---|
| Long-run coefficients | | | | | |
| lnTAX | | -0.043 | -0.333* | 0.084*** | -1.297 |
| lnGINI | 0.794*** | | 1.564*** | -0.074 | -4.194 |
| lnTRADE | -0.212** | 0.089 | | 0.051*** | 5.299 |
| lnURBAN | 0.913** | -1.525*** | 1.772*** | | 3.962 |
| lnPCGDP | -0.009 | 0.959*** | -0.676*** | -0.054** | |
| ECT | -0.229*** | 0.241** | -0.149** | 0.135 | -0.003 |
| Short-run coefficients | | | | | |
| lnTAX | | -0.043 | 0.282 | 0.004 | 0.081** |
| lnGINI | 2.543 | | 0.275 | 0.047* | 0.149 |
| lnTRADE | -0.049 | -0.062* | | -0.002 | 0.002 |
| lnURBAN | -112.324 | 2.247 | 45.787 | | -0.232 |
| lnPCGDP | 1.319*** | 0.417** | 0.547 | -0.067 | |

## 4. FINDINGS

On the assumption of long-run homogeneity, we bypassed the restrictive panel cointegration tests. In our case, cointegration has been confirmed by the statistical significance of the long-run coefficients of the PMG estimation and the ECT terms presented in Table 3. The PMG restricts the long-run equilibrium to be homogeneous across countries while allowing for heterogeneity in the short-run relationship (Pesaran et al., 1999). The PMG estimator has been noted to be robust, and we have used the Hausman test to verify the appropriateness of using the PMG estimator for our postulated models. For examining the short and long-run relationships between inequalities (lnGINI) and taxes (lnTAX) and other variables, the panel ARDL – initiated by Pesaran and Smith (1995) and Pesaran et al. (1999) – is capable of handling the underlying regressors that exhibit a mixture of I(0) and I(1) while none of the variables is I(2) (see Pesaran and Shin, 1999) with a time span of over 20 years. From Table 3, our main findings are as follows:

- Model 1 and Model 3 are the only ones that display long-term cointegration since the ECT term is negative and statistically significant, while the long-term coefficients are also statistically significant.
- The results for Model 1 establish that causality runs from inequality (lnGINI) to taxes (lnTAX) as taxes and inequality bear a positive relationship. In the short-run, the per capita income (lnPCGDP) has a positive impact on taxes, but not in the long-run.

- Model 2 shows that taxes do not have a long-term impact on inequality. In other words, there is no evidence of causality running from taxes to inequality.
- Model 3 shows that the internationalisation of the economies of the chosen nations – measured by lnTRADE – has been driven by both taxes (-) and inequalities (+), while there seems to be evidence of mutual causality between taxes (lnTAX) and internationalisation (lnTRADE). Thus, further empirical development will be necessary to fully unravel the comprehensive interrelationships between taxes (lnTAX) and internationalisation (lnTRADE).
- Model 4 and Model 5 show that neither urbanisation nor development (per capita GDP) bears a long-run relationship with income inequality and other control variables.

Model 1 shows that the elasticity of taxes (as a percentage of GDP) with respect to income Gini is 0.8 and statistically significant at the 1% level. In other words, if the income Gini increases (declines) by 1%, tax revenues (as a percentage of GDP) increase (decrease) by 0.8% in the long run. As the ECT term shows, deviations from the equilibrium get corrected by 22.9% (-0.229) annually for the chosen countries. Model 1 confirms the prediction of the median voter hypothesis on the consequences of income distribution: greater inequality is associated with a larger tax-GDP ratio because of the greater redistribution that is sought by the median voter when income distribution is less equal (see Milanovic, 2000, 2003). We did not find an impact of per capita GDP (lnPCGDP), as a measure of overall development, to exert any influence upon taxes.[7]

Internationalisation (lnTRADE) is known to have two mutually opposing impacts on taxes, and we find that lnTRADE has a negative, statistically significant, long-term effect on lnTAX for the countries under consideration.[8]

We also note that urbanisation (lnURBAN) has a positive – and statistically significant – long-term impact on taxes.[9] The elasticity of taxes (as a percentage of GDP) to urbanisation is positive and statistically significant at a 5% level. An increase (decrease) in urbanisation by 1% leads to an increase (decrease) in the tax-GDP ratio by 0.913%.

## 5. CONCLUSION

The existing literature has extensively examined the impacts of taxes on inequalities within and between nations. It is only recently that the empirical consequences of

[7] In the existing literature, the effect of per capita GDP on taxes has been ambiguous: on the one hand, it is expected to have a positive effect because as a country experiences a higher level of development, the formalisation and the competitiveness of the economy expand with expanding possibilities for taxes. On the other hand, an open economy reduces tariffs and trade barriers and this fact can have negative effects on tax collection (Baunsgaard & Keen, 2010). Depending on the relative strengths of these two effects, the overall effect of PCGDP on TAX is determined. So for the ACI economies, the two opposing forces seem to nullify each other.

[8] Foreign trade and investment are known to boost taxes by improving competitiveness and the formalisation of an economy (see Cassou, 1997; UNCTAD, 2000; Martín-Mayoral & Uribe, 2010; Gugler & Brunner, 2007). With increased international trade, the tax bases also shift from the domestic economies leading to lower taxes, *ceteris paribus* (see Baunsgaard & Keen, 2010).

[9] It is well-recognised in the existing literature that urbanisation is driven by increasing roles for the industrial and services sectors. Not only do the industrial and services sectors have a large tax base – *vis-à-vis* the agricultural sector – it is easier to tax industrial enterprises than agricultural enterprises. Thus increased urbanisation is expected to increase taxes (see Eltony, 2002).

inequalities have been explored for the tax-GDP ratios of the OECD economies: Islam et al. (2018) – contradicting the theoretical predictions of the median voter models – showed that inequalities significantly lowered the income tax to GDP ratios of OECD countries over a very long horizon spanning from 1870 to 2011. One of the apparent weaknesses of the work of Islam et al. (2018) is their neglect of the (potential) reverse causality coming from taxes to inequalities, which has been a well-received doctrine in the extant literature. In this work, we establish that there is no evidence of the reverse causality from taxes to inequalities for the ACI economies. The main methodological weakness in the work of Islam et al. (2018) is the known inadequacy of their panel models to handle non-stationarity and autocorrelation of variables over a long horizon (see Bruckner & Ciconne, 2011; Ciconne, 2011, 2013). Hence we chose the (panel) ARDL methodology, being robust to autocorrelation, non-stationarity and mild endogeneity, that can simultaneously handle both stationary and nonstationary variables (see Gangopadhyay & Nilakantan, 2018).

By applying a robust model based on the panel ARDL methodology – for another set of countries where inequalities and taxes play a significant role – we are able to establish that there is a long-run relationship between income inequalities and taxes with the causality running only from inequalities to taxes. With no evidence of causality from taxes to inequalities for ACI economies, our results are credible. However, contrary to the findings of Islam et al. (2018), our results fully support the theoretical predictions of the median voter models that argue that increases in inequalities will increase taxes. We found that this elasticity of taxes with respect to income Gini is inelastic (0.8).

Our results suggest that growing income inequality can incentivise governments towards favouring populist policies in the ACI economies: for instance, as income inequality rises, lower-income voters demand higher taxes and stricter regulation (see Persson & Tabellini, 1994). Moreover, governments are also aware of a potential erosion of trust, caused by rising inequality, which can increase tax non-compliance among citizens (see Tran-Nam & Le, 2022). A loss of trust can trigger serious political crises, or even political instability, as argued by Keefer and Knack (2002). Hence, rational governments will have an incentive to raise taxes in response to rising inequality to control tax non-compliance (Tran-Nam & Le, 2022). Nonetheless, as Harberger (2003) stressed, any attempt to impose unduly redistributive taxes might backfire.

Thus, our study has the following three policy recommendations for the ACI region. First, as Huang, Morgan and Yoshino (2019, p. 1) note, similarly to the findings of Zhuang, Kanbur and Rhee (2014, pp. 38-39), a large proportion of inequality in Asia stems from barriers to education and problems with human capital formation. Increased taxes can finance public spending for improving access to education, augmenting human capital and skills of the weaker sections of society. Improved access to education can, in turn, help lower long-run income inequality and thereby lower the needs for future taxation. Short-term policy implications might include transfers to low-income families for improving their health and educational outcomes, which can reduce future inequality and, in turn, the burden of taxation. Secondly, classical developmental theories posit urbanisation as a critical means for reshaping emerging economies – burdened with a large stock of surplus labour in the rural sector. Urbanisation can help narrow the urban-rural gap and, thereby, reduce income inequality (Wan & Zhuang, 2015)[10] and, hence, lower the needs for larger taxes. Finally, governments in the ACI region must promote policies that create equal access to public goods and services, alleviate corruption to enhance institutional quality and governance and reduce inequality to lower future taxes for promoting inclusive (economic) growth (Dollard & Miller, 1950, ch. 3).

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[10] However, as hypothesised by Kuznets (1955), it may take a longer timeframe before urbanisation mitigates income inequality. Hence, investments in developing communication and transport infrastructure in rural and inland areas, increasing their connectivity with urban economic hubs, could be a more efficient and immediate remedy.

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