Government Intervention, Institutional Quality, and Income Inequality: Evidence from Asia and the Pacific, 1988–2014

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We examine the linear and nonlinear long-run relationship between public expenditure and institutional quality, and income inequality in Asia and the Pacific. By applying panel cointegration methods using a dataset from 1988 to 2014, our main findings suggest that public expenditure and institutional quality have negative long-run, steady-state effects on income inequality in Asia and the Pacific. The effect of institutional quality has only a one-way Granger causality link to income inequality. The existence of a nonlinear relationship between public expenditure and institutional factors linked to income inequality is also found. It implies that, at the early stage of institutional development, a country whose economy has experienced higher public expenditure generates rising income inequality; then, in the long run, when the country improves its institutional quality, higher public expenditure results in lower income inequality.

Keywords: Asia and the Pacific, income inequality, institutional quality, public expenditure
JEL codes: D63, E02, H53

I. Introduction

During the last few decades, the Asia and Pacific economies have achieved impressive economic development compared to the global average; however, the region is lagging with respect to rising economic inequality (Alvaredo et al. 2018, United Nations 2018). Hollywood's romantic dramedy, *Crazy Rich Asians*, and the black comedy thriller from the Republic of Korea, *Parasite*, are recent pieces of

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Asian Development Review, vol. 38, no. 1, pp. 176–206
https://doi.org/10.1162/adev_a_00162

social evidence that explain how far the highest income groups are from the rest of society in some Asian countries. Piketty (2014) argued that it would be a mistake to underestimate the importance of film and literature, 19th century novels especially, which are full of detailed information about the relative wealth and living standards of different social groups, the deep structure of inequality and the way it is justified, and its impact on individual lives. According to the last updated data from Credit Suisse's global wealth report and Oxfam, the number of super-rich (millionaires and billionaires) in the Asia and Pacific region—comprising Australia; Hong Kong, China; India; Indonesia; Japan; the People's Republic of China (PRC); the Republic of Korea; Taipei,China; Thailand; and Viet Nam; among other economies—has surpassed that of North America and Europe. In another sign of rising inequality, Asia and the Pacific's income Gini coefficient increased from 0.37 to 0.48 between 1990 and 2014, while the gap in wealth equality is even wider; in addition, the Asia and Pacific region is also the home of nearly two-thirds of the world's working poor (Costa 2018).

During the early stage of its economic development a half-century ago, the Asia and Pacific region was widely known as a place where there were countless internal conflicts, political instability, civil wars, and widespread poverty. During that period, the region's governments focused on offering basic needs and instituting a minimum degree of social security, as well as securing the rule of law for their citizens (Sobrado et al. 2014). As in most advanced economies, as a society prospers, people's expectations become more demanding in terms of access to better government services—including the rule of law, accountability, transparency—and a simultaneous improvement of welfare and income redistribution. As reported by the most recent findings, rising income and wealth inequality is considered among today's biggest challenges for governments around the world, and the Asia and Pacific economies are no exception (Stiglitz 2012, Piketty 2014). In the long run, lessons from history show that an unequal society can lead to global disaster, like what Europeans faced twice in the 20th century. For that reason, this issue does matter above all.

In the modern era of globalization, many aspects—from an economy's factor endowments, trade openness, financial deepening, geography, and institutions, to its historical trajectory and technological changes—have been detected to explain inequality (see, for example, Kanbur and Zhuang 2013 and Hartmann et al. 2017). Therefore, to address the inequality issue, it is, without doubt, a matter of controversy and complexity. Besides a variety of redistributive policies, government intervention through public expenditure is normally prioritized in developing economies. This is because taxation mechanisms are viewed as less effective and less efficient, thanks to the small size of tax revenues and low quality of governance and institutions. Nyblade and Reed (2008) suggested that public expenditure is able to function well and promote a more equal society only if the institutional quality in some context (e.g., low level of corruption and high political competition) is allowed to do it. In Asia and the Pacific, however, the institutional framework has been seen to progress slowly like a crab, moving forward then backward from time to time. Accordingly, a closer look at this specific issue is required.

This paper aims to investigate the linear and nonlinear long-run relationship between public expenditure and institutional quality, and inequality in Asia and the Pacific. To realize this, we use a dataset covering eight countries in Asia and the Pacific—Australia, India, Japan, Malaysia, the PRC, the Republic of Korea, Singapore, and Thailand—from 1998 to 2014.

The contribution of this paper to the literature is as follows. First of all, it gives new empirical findings on the long-run impact of public expenditure and institutional quality on income inequality, which previously has been extensively studied only in the short run and medium run. Combining the strength of panel fully modified ordinary least squares (FMOLS) and panel dynamic ordinary least squares (DOLS) with Granger causality tests, our approach can examine simultaneously the effect in Asia and the Pacific of public expenditure and institutional quality on income inequality, and the effect of income inequality on public expenditure and institutional quality. Secondly, the nonlinear panel cointegration models are designed to investigate the nonlinear relationship of public expenditure and institutional quality on income inequality across the sample countries. This may be one of the pioneering theses in applying the nonlinear long-run panel models to estimate a hypothesis, particularly in the Asia and Pacific region. Thirdly, it is applied to a new measurement. We used the World Inequality Database (WID.world), first developed by Piketty and Zucman (2014), which we found a more available dataset for the Asia and Pacific countries in this study. This new measurement provides further insight into the thinking of inequality. Our dataset covers at least 26 years (1988–2014) for almost all countries.

Finally, this paper focuses on specific countries in Asia and the Pacific. It is complementary to the existing literature, which we found focused mainly on the advanced economies in Europe and the United States (US). However, when the global economy changes, the standard model of economic thoughts should also change. This matters for the Asia and Pacific economies, which might not follow a similar pattern of development as that of other countries. As Robert Solow explained, there is no economic theory of everything (Todaro and Smith 2017). In the 21st century, when the center of gravity of the world economy has shifted decisively from the Atlantic to the Pacific Ocean, everything that happens in these two regions will attract very strong public attention. Since the Asia and Pacific region has not been empirically studied as extensively as Europe and the US, this study seeks to provide some insights in light of the controversial findings from some previous studies.

The rest of the paper is organized as follows. Section II considers the literature review. Section III explains the empirical methodology and data, and then provides the testing results from panel unit root tests and panel cointegration tests. Section IV looks at the overall regression results and presents a discussion of the results. Section V discusses the robustness checks. The final section gives the concluding remarks.

II. Literature Review

A. The Effect of Public Expenditure on Inequality

Besides taxes, government intervention may help to reduce inequality by redistributing resources through public expenditure (Doerrenberga and Peichla 2014). In this perspective, the government might implement a wide range of mechanisms through transfers involving education, health, social insurance, housing, infrastructure, public investment, and other welfare programs (Gruber 2013). Progressive taxation is a common policy measure for reducing inequality, not only individual but also corporate taxation may impact on individual income and wealth (e.g., Hazak 2009). Another public finance channel used to address inequality is (progressive) government spending, and there are many theories and pieces of evidence to suggest that certain sorts of public spending policies are likely to promote a more equal society.

For instance, the human capital theory argues that investment in further education tends to increase a person's stock of skills and productivity (Gruber 2013). Thus, education may promote a better outcome in society. In some particular contexts, government intervention—for instance, providing subsidies to low-income families for early-education investments to mitigate young parents' budgetary concerns—could have a significant role to play in providing equal access to education, which consequently decreases income inequality and increases intergenerational mobility (Juan and Muyuan 2016).

Empirically, although higher education has expanded significantly on a global scale, it is suggested that we are living in a less equal world. One important perspective is the contribution of human capital and investments in research and development to growth along with convergence (Männasoo, Hein, and Ruubel 2018), but this alone does not guarantee that the benefits of increasing knowledge intensity are equally or fairly distributed. In the Asia and Pacific region, we observe that participation in higher education is increasing rapidly in most countries but, at the same time, social mobility lags behind the development of higher education (Marginson 2018). We also find rising wealth and income inequality in advanced English-speaking countries even as they have many of the top universities in the world (Piketty 2014). However, given that education from primary to tertiary is free or almost free in European welfare countries—such as France, Germany, and Scandinavian countries like Denmark, Finland, and Sweden—we observe that they are more equal societies in terms of wealth and income distribution. According to Marginson (2018), many higher education systems are more vertically stratified, with a larger stretch in status and resources between top universities and other higher education institutions. Elite universities tend to be dominated by students from advantaged backgrounds, blocking the potential for greater social mobility, though their social composition varies from case to case.

In this regard, the effectiveness of distributive public policies would be necessitated to go along with a particular assumption or hypothesis. There is much evidence to argue that public expenditures that target the lower and lower-middle social classes, which comprise the majority of the population, would produce a more equal distribution of outcomes. This supports the idea that public policies need to be involved in providing basic health insurance, compulsory education (primary and secondary), unemployment insurance, housing subsidies, and public infrastructure (Gruber 2013). Considering policy and implementation, it becomes not only complex but also complicated. Some studies suggest that although public policies may be designed to target the most vulnerable or the most needy citizens at the early stage, the benefits might end up going to the middle or elite social classes. It might be due to government failure, corruption, or low quality of good governance and institutions. This evidence can be found in many low-income and middle-income developing countries (Anderson et al. 2017).

Another consideration is to view things in both the short and long run. Let us suppose we are living in a world where we have an equal degree of good governance and institutions so that the government can function at the highest efficiency (lowest rate of corruption or least possible government failure). In this case, even though public expenditure tends to reduce inequality in the short run, it does not guarantee that inequality is less likely to worsen in the long run. Bourguignon (2004) states that too many income transfers, as opposed to transfers of wealth, can lower the expected return from acquiring physical and human capital. They might distort the economy and reduce savings and investment, and therefore the rate of growth. According to Lee (2013), greater government income transfers may reduce people's incentive to work for themselves, and then the whole economy becomes less dynamic. Consequently, it could generate a possible economic recession in the long run. If this hypothesis were right, citizens from the lower and middle classes would find themselves struggling more than the higher social classes during the crisis. Thus, inequality might be rising subsequently.

B. The Effect of Institutional Quality on Inequality

According to Zhuang, de Dios, and Lagman-Martin (2010), we can associate institutional quality with inequality in two different ways: (i) political institutions and democracy, and (ii) corruption. On the one hand, in relation to political factors, it has been suggested that more equal income distribution would be better promoted in a democratic society with more political rights. When political rights to vote are extended to the majority of the population, the amount of redistribution is decided by the median voter and this determines, directly or indirectly, the rate of growth of the economy (Bourguignon 2004). However, it has failed to be verified empirically in some cases where countries with a higher score of democracy are not necessarily reducing inequality. It is subject to the fact that the political system alone cannot explain inequality. For example, despite having a lower score in democracy or restrictive political rights, income distribution in many countries—such as East European countries, the Republic of Korea, and Singapore—was relatively equal as long as their respective societies functioned with a special political ideology. Moreover, democracy is more likely to reduce inequality in countries with a parliamentary rather than a presidential system (Zhuang, de Dios, and Lagman-Martin 2010).

Corruption, on the other hand, tends to increase income inequality for the reason that it can lead to tax evasion, less effective administration, lower progressive taxes, less effective public expenditure, and lower investment. The problem would potentially create political, economic, and social systems that favor only the rich and hurt the poor (Pedauga, Pedauga, and Delgado-Márquez 2017). In contrast, some argue that corruption can lead to less inequality if the social benefit from corrupted activities is greater than the social damage. Another recent study has found that corruption tends to be associated with lower inequality in less developed countries due to the existence of the informal sector in many developing countries (Andres and Ramlogan-Dobson 2011). In the analysis of more disaggregated data, Nyblade and Reed (2008) have linked corruption to inequality in two contexts: (i) political competition and (ii) voting. The first involves corrupt actions to gain personal benefits by the elites in society, which would increase inequality. The second, however, involves buying votes by using, for instance, public budgets to reach the mass of the population. This tends to decrease inequality because at least the money goes to the poor people.

In addition, there are multiple channels through which institutions may impact inequality. For example, various social norms may propagate inequality among different population groups (e.g., some ethnic groups, minorities, and females), and rent-seeking opportunities may foster inequality and financial constraints (e.g., Männasoo, Maripuu, and Hazak 2018) that often have an institutional background that may affect different types of individuals and companies differently.

Linking together government intervention through public expenditure and inequality in the context of diverse institutional quality, we might consequently presume that the distributive effect of public expenditure tends to reduce inequality, given that an economy has high-quality governance and institutions. If this hypothesis is not completely right, the policies would not be implemented effectively. Alternatively, in cases of low institutional quality or high corruption, public intervention tends to increase income inequality because it would lead to tax evasion, less effective administration, lower progressive taxes, less effective public expenditure, and lower investment. However, it is likely to promote more quality outcomes only if the existence of social benefits, provided by public intervention, is linked to the mass of the population (i.e., the poor), such as social assistance or gift giving during election. This hypothesis does not take into account its effects in the medium and long run, which are complex by nature.

III. Empirical Methodology

A. Data

We collected data from various sources from 1988 to 2014 in the following countries: Australia, India, Japan, Malaysia, the PRC, the Republic of Korea, Singapore, and Thailand.

We used the pretax top 1% income share of the population to measure income inequality. The data were taken from the World Inequality Database (WID.world), first developed by Piketty and Zucman (2014). Our dataset is available for at least 26 years (1988–2014) for most countries, except for the Republic of Korea (1995–2014), Thailand (2000–2014), and Malaysia (a total of 13 years with missing values and another 13 years with data between 1988 and 2014). To deal with missing values, we applied the cubic-spline interpolation methods as explained by McKinley and Levine (1998); Fichtenbaum and Shahidi (1988); and Bishop, Chiou, and Formby (1994). The rationale for using this indicator follows the theses of Malinen (2016), who presented arguments linking income inequality to credit cycles, and Leigh (2007), who argued that there is a strong and significant relationship between top income shares and broader inequality measures, such as the Gini coefficient. According to Malinen (2016), the top 1% income share measures the share of national income concentrated in the hands of the highest percentile of income earners. As gross domestic product (GDP) is, in practice, the national income of a country, the share of total income received by the top 1% of earners can also be presented as $\frac{\textit{income of the top } 1\%}{GDP}$. However, to avoid the bias that the top 1% income share cannot capture the full picture of the effect of public spending and institutional quality in promoting economic opportunity for the poor and the middle class, we also employed version 8.2 of the Standardized World Income Inequality Database (SWIID) of Solt (2019) for robustness checks. It is the estimate of the Gini index of inequality in equivalized (square root scale) household disposable (posttax, posttransfer) income, using the Luxembourg Income Study data as the standard. The SWIID dataset is available for nearly 100% of our eight sample countries in Asia and the Pacific.

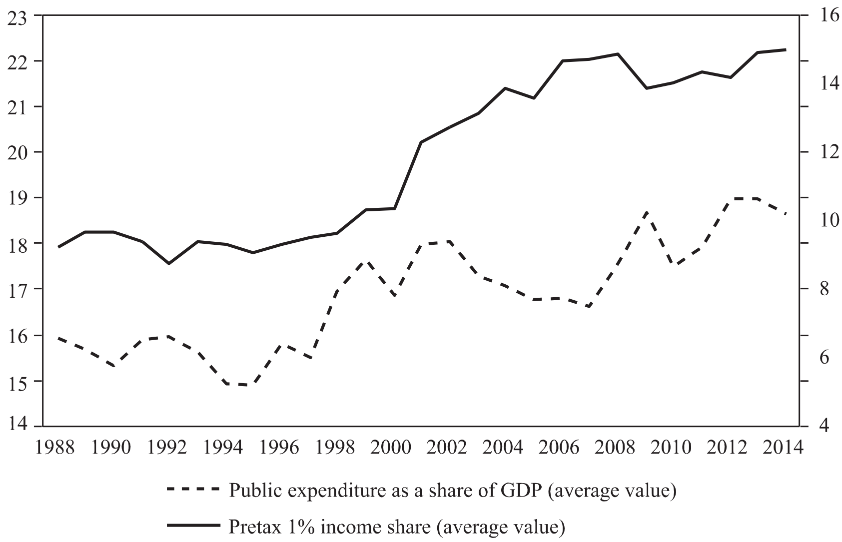
To make our estimation comparably reasonable, we used public expenditure (share of GDP): $\frac{\textit{Public expenditure}}{GDP}$. Public expenditure comprises cash payments for the operating activities of the government in providing goods and services. It includes compensation of employees (e.g., wages and salaries); interest and subsidies; grants;

social benefits; and other expenses such as rent and dividends (based on World Bank definitions). To investigate the role of institutional quality, we used the average value of the Worldwide Governance Indicators (WGI), which are found in the empirical works of Zhuang, de Dios, and Lagman-Martin (2010); Kaufmann, Kraay, and Mastruzzi (2010); and Wong (2017). The WGI consists of six broad dimensions of governance: (i) voice and accountability, (ii) political stability and absence of violence and terrorism, (iii) government effectiveness, (iv) regulatory quality, (v) rule of law, and (vi) control of corruption. The estimate of governance performance in standard normal units ranges from approximately −2.5 (weak) to 2.5 (strong). The Asia and Pacific countries that are defined as having strong institutional quality have an average WGI value that is "bigger or equal to zero"; otherwise, they are defined as having weak institutional quality. Therefore, Australia, Japan, Malaysia, the Republic of Korea, and Singapore are in a group of countries with strong institutional quality. India, the PRC, and Thailand are in a group of countries with weak institutional quality.

In addition to explanatory indicators, we added several major aggregated variables as additional control variables (Appendix Table A.1). The country's openness is theoretically linked to income distribution (see, for example, Heckscher 1919, Ohlin 1933, Samuelson 1953, and Melitz and Redding 2015). The sum of imports and exports is used to measure trade openness (see, for example, Cameron 1978, Rojas-Vallejos and Turnovsky 2017, and Wong 2017). The level of development is also linked to inequality. The general effect of GDP per capita on income inequality is explained by the well-known inverted-U hypothesis developed by Simon Kuznets: an increase in GDP per capita will increase overall economic welfare and income disparity. Following the process of economic development, inequality will increase during the first stage; after it arrives at the peak, inequality will decrease (Kuznets 1955). Changes affecting labor supply and labor demand can also shift income inequality. Changes in population, measured by the annual percentage growth in population, affect changes in labor supply and demand, which affect wages in the labor market. An increase in population is expected to increase income inequality if the unemployment rate increases (see, for example, Asteriou, Dimelis, and Moudatsou 2014; Rojas-Vallejos and Turnovsky 2017; Wong 2017). Oil rents (as a share of GDP) are used to account for resource-rich regimes that can afford to gain legitimacy by redistributing revenue (Ross 2001, Wong 2017). Taxation, in addition to public spending, may also target to improve the overall economic well-being of a whole population, especially the poor. The effect on income distribution depends on how the government targets specific population groups through social protection, education, and health, among others (Selowsky 1979, Younger 1999).

The addition of listed control variables into the model may impair the identification of individual coefficients in the presence of high multicollinearity. As shown in Appendix Tables A.2 and A.3, the variance inflation factor and the

Figure 1. **Average Value of Top 1% Income Share and Public Expenditure, 1988–2014 (% of GDP)**



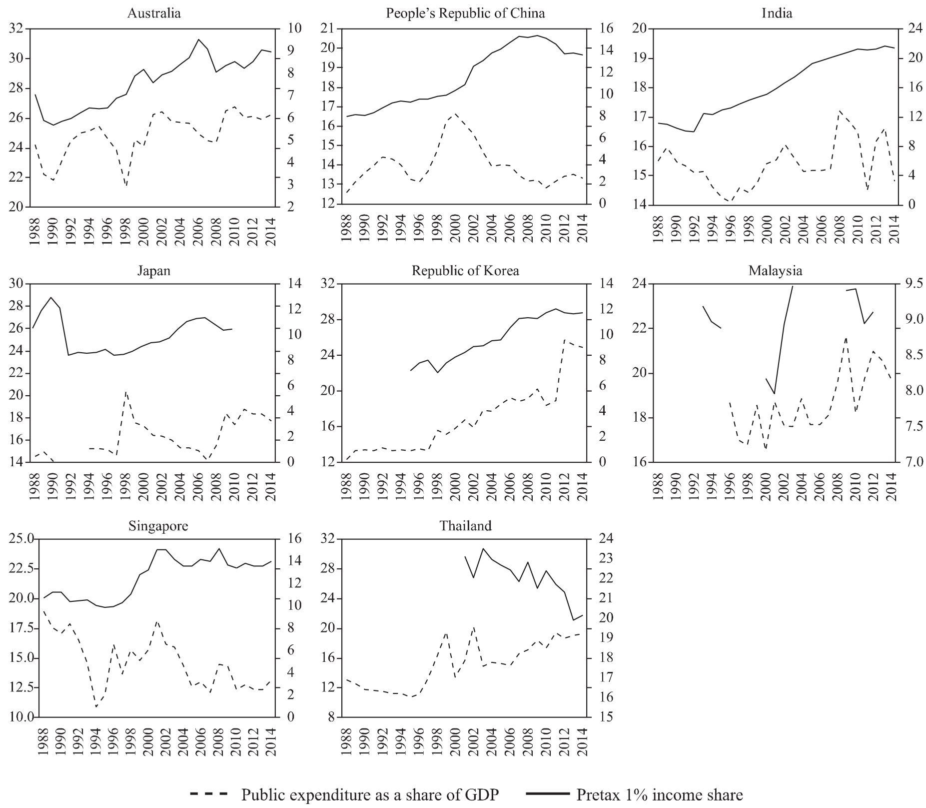
GDP = gross domestic product.
Source: Authors' calculations using the World Inequality Database. 1988–2014. WID.world (accessed December 3, 2018).

pairwise correlations among explanatory variables, however, did not reveal any severe multicollinearity. The variance inflation factor was 4.04, well below the critical level of 10. The pairwise correlation estimates confirmed that correlations between variables were well below the critical levels.

The list of countries divided by regions, income status, and institutional status along with descriptive statistics of the Asia and Pacific countries are reported in Appendix Tables A.4 and A.5. India is the only lower-middle-income country from South Asia; the others are all from East Asia and the Pacific. Malaysia is the only country with upper-middle-income status, but it is classified in the same group as high-income countries with strong institutional quality. Figure 1 shows that public expenditure followed a rising trend from the end of the 1980s to 2014. The top 1% income share, on the other hand, evolves in a stable trend then starts increasing from the early 1990s; overall, it also shows a rising trend from 1988 to 2014. The results give some evidence to the extent of the "trending hypotheses," indicating a possible long-term correlation between the variables. However, that can be a reverse causality (i.e., public expenditure explains the top 1% income share and vice versa).

Figures 2 shows a different pattern in each country. Australia, India, Malaysia, and the Republic of Korea, show both a rising trend for public expenditure and the top 1% income share. The PRC and Singapore show only a rising trend for the top 1% income share and a nearly stable trend of public expenditure. This is the

Figure 2. **Top 1% Income Share and Public Expenditure, 1988–2014 (% of GDP)**



GDP = gross domestic product.
Sources: Authors' calculations using the World Inequality Database. 1988–2014. WID.world (accessed December 3, 2018); and World Bank. 1988–2014. World Development Indicators. https://databank.worldbank.org/source/world-development-indicators (accessed December 3, 2018).

opposite for Japan, where public expenditure is rising but the top 1% income share is likely to become stable in the long run. Thailand shows a very different trend than other countries, as there is a reverse trend between the variables.

To assess the long-run equilibrium association between the variables, we performed several tests, including the panel unit root test, panel cointegration test, and cointegration regression estimation.

B. Panel Unit Root Tests

We used three types of panel unit root tests. The first follows unit root assuming individual unit root process, including the Im–Pesaran–Shin test (2003), Augmented Dickey–Fuller (ADF)–Fisher test by Maddala and Wu (1999), and Phillips–Perron (PP)–Fisher test by Choi (2001). The second follows unit root assuming common unit root process, including the Levin–Lin–Chu test (2002). The third allows for homoscedastic error processes across the panel, including the tests of the Hadri *Z*-stat and heteroscedastic consistent *Z*-stat by Hadri (2000).

The panel unit root tests, in this paper, are based on the following regression equation:

$$\Delta y_{it} = \rho_i y_{i,t-1} + \alpha_i + \eta_i t + \theta_t + \varepsilon_{it} \tag{1}$$

where α_i are individual constants; $\eta_i t$ are individual time effects, and θ_t are the common time effects. The null hypothesis of individual unit root process is that the panel data has unit root, $H_0 : \rho_i = 0 \; \forall\, i$, (i.e., the series in $I[0]$ are nonstationary).

The alternative hypothesis is as follows:

$$H_1 : \rho_i < 0, i = 1, 2, \ldots, N_1, \rho_i = 0, i = N_1 + 1, N_2 + 2, \ldots, N.$$

The same principle is applied for the Levin–Lin–Chu test, assuming common unit root process. However, the null hypothesis of the tests of Hadri *Z*-stat and heteroscedastic consistent *Z*-stat is that panel data has no unit root (the process is stationary), and the alternative hypothesis is that the panel data has unit root (the process is nonstationary).

According to the results of the panel unit root tests from Table 1, six among six tests for the top 1% income share, public expenditure (% of GDP), trade (% of GDP), per capita GDP at purchasing power parity (PPP), and tax revenue (% of GDP), and four among six tests for population growth (annual %) emphasize that the majority of tests become nonstationary at level; then, the series become stationary after first difference, $I(1)$.

C. Panel Cointegration Testing

To estimate the panel cointegration model, the series of our variables must be nonstationary at level, $I(0)$, but become stationary at first difference, $I(1)$. This condition is confirmed in our model. We applied, therefore, the Pedroni residual cointegration test (Pedroni 1999, 2004); Kao residual cointegration test (Kao 1999); and Fisher–Johansen cointegration test (Maddala and Wu 1999).

The models for testing panel cointegration between income inequality and public expenditure are structured as follows:

$$inequality_{it} = \alpha_i + \beta_i expense_{it} + \eta_i t + \theta_t + \varepsilon_{it} \tag{2}$$

where $(1, -\beta_i)$ are country-specific cointegrating vectors, α_i are individual constants, $\eta_i t$ are individual time effects, and θ_t are the common time effects. The null hypothesis is that $H_0 : \beta_i = 1 \; \forall\, i$ (i.e., there is no cointegration).

Table 1. **Panel Unit Root Tests**

	Individual Unit Roots			Common Unit Roots	Heteroscedastic	
Series	**IPS**	**ADF-Fisher**	**PP-Fisher**	**LLC**	**Hadri**	**Hetero Con Z-stat**
Top 1%	-0.732^{***}	20.845^{***}	21.062^{***}	0.358^{***}	3.807^{***}	3.318^{***}
income share	(0.232)	(0.185)	(0.176)	(0.640)	(0.000)	(0.001)
Public expenses	-0.207^{***}	21.581^{***}	17.752^{***}	-0.826^{***}	2.836^{***}	4.100^{***}
(% of GDP)	(0.418)	(0.157)	(0.339)	(0.205)	(0.002)	(0.000)
Institutional	−1.704	32.045	24.021^{***}	−2.189	2.410^{***}	4.792^{***}
quality	(0.044)	(0.010)	(0.089)	(0.014)	(0.008)	(0.000)
Trade	0.169^{***}	12.266^{***}	11.483^{***}	-0.589^{***}	6.131^{***}	5.534^{***}
(% of GDP)	(0.567)	(0.726)	(0.779)	(0.278)	(0.000)	(0.000)
Per capita GDP	-0.834^{***}	22.293^{***}	19.556^{***}	-1.265^{***}	3.514^{***}	2.763^{***}
(PPP)	(0.202)	(0.134)	(0.241)	(0.103)	(0.000)	(0.003)
Population growth	-1.212^{***}	32.619	19.089^{***}	3.011^{***}	0.753	4.295^{***}
(annual %)	(0.113)	(0.008)	(0.264)	(0.999)	(0.226)	(0.000)
Oil rents	−2.142	24.299	24.906	−4.178	3.280^{***}	4.670^{***}
(% of GDP)	(0.016)	(0.042)	(0.036)	(0.000)	(0.001)	(0.000)
Tax revenue	-0.053^{***}	22.1816^{***}	16.556^{***}	-1.3513^{***}	4.030^{***}	4.462^{***}
(% of GDP)	(0.300)	(0.137)	(0.415)	(0.088)	(0.000)	(0.000)

ADF = Augmented Dickey–Fuller, GDP = gross domestic product, IPS = Im–Pesaran–Shin, LLC = Levin–Lin–Chu, PP = Phillips–Perron, PPP = purchasing power parity.
Notes: All tests are taken using automatic selection of maximum lags; automatic lag length selection based on Schwarz information criterion; Newey–West automatic bandwidth selection and Bartlett kernel; assumed asymptotic normality and individual effects; and individual linear trends, except for public expenditure (% of GDP), which we include only for individual effects. *** emphasizes that the process is nonstationary at level, then becomes stationary at level $I(1)$ after we reject or do not reject the null hypothesis.
Sources: Authors' calculations using the World Inequality Database. 1988–2014. WID.world (accessed December 3, 2018); and World Bank. 1988–2014. World Development Indicators. https://databank.worldbank.org/source/world-development-indicators (accessed December 3, 2018).

Tables 2, 3, and 4 report the results from the Pedroni residual cointegration test, Kao residual cointegration test, and the Fisher–Johansen cointegration test with a dataset from 1988 to 2014. According to the estimated results from various panel cointegration tests, indicating that most of the statistics have a p-value of less than the 1% and 5% level of significance, the null hypothesis of no cointegration can be rejected. Therefore, we conclude that there is a high possibility of a long-run equilibrium relationship between public expenditure and income inequality in the Asia and Pacific countries.

IV. Results and Discussions

A. Estimating a Cointegrating Regression

To obtain the long-run coefficients between the variables of interest, we took into account two different but complementary estimators. First, we estimated with the FMOLS by Phillips and Hansen (1990). Second, we estimated with the DOLS by Stock and Watson (1993), and Mark and Sul (2003).

Table 2. **Pedroni Residual Cointegration Test**

Within dimension				
			Weighted	
	Statistic	**Prob**	**Statistic**	**Prob**
Panel v-Statistic	−1.278	0.899	−0.562	0.713
Panel rho-Statistic	−3.410***	0.000	−2.175**	0.015
Panel PP-Statistic	−4.870***	0.000	−3.073***	0.001
Panel ADF-Statistic	−5.494***	0.000	−3.683***	0.000
Between dimension				
	Statistic	**Prob**		
Group rho-Statistic	1.120	0.869		
Group PP-Statistic	0.131	0.552		
Group ADF-Statistic	−0.366	0.357		
Countries	8			
Observation	216			

ADF = Augmented Dickey–Fuller, PP = Phillips–Perron.
Notes: Null hypothesis = no cointegration. The tests were estimated with the following assumptions: trend assumption (deterministic intercept and trend), automatic lag length selection based on Schwarz information criterion with lags from 0 to 5, and Newey–West automatic bandwidth selection and Bartlett kernel. ***1% level of significance, **5% level of significance, *10% level of significance.
Sources: Authors' calculations using the World Inequality Database. 1988–2014. WID.world (accessed December 3, 2018); and World Bank. 1988–2014. World Development Indicators. https://databank.worldbank.org/source/world-development-indicators (accessed December 3, 2018).

Table 3. **Kao Residual Cointegration Test**

	***t*-Statistic**	**Prob**
ADF	2.274**	0.012
Residual variance	1.414	
HAC variance	0.792	
RESID(−1)	−0.114**	0.012
D(RESID[−1])	−0.280***	0.001
Observations	216	

ADF = Augmented Dickey–Fuller, HAC = heteroscedasticity- and autocorrelation-consistent, RESID = residual.
Notes: Null hypothesis = no cointegration. The tests were estimated with the following assumptions: trend assumption (no deterministic trend), automatic lag length selection based on Schwarz information criterion with a max lag of 1, and Newey–West automatic bandwidth selection and Bartlett kernel. ***1% level of significance, **5% level of significance, *10% level of significance.
Sources: Authors' calculations using the World Inequality Database. 1988–2014. WID.world (accessed December 3, 2018); and World Bank. 1988–2014. World Development Indicators. https://databank.worldbank.org/source/world-development-indicators (accessed December 3, 2018).

Table 4. **Fisher–Johansen Cointegration Tests**

Hypothesized No. of CE(s)	**Fisher Stat[a] (from trace test)**	**Prob**	**Fisher Stat[a] (from max-eigen test)**	**Prob**
None	62.12***	0.000	52.69***	0.000
At most 1	37.45***	0.001	37.45***	0.001
Observations	216			

CE = cointegrating equation.
[a]Probabilities are computed using asymptotic chi-square distribution.
Notes: Null hypothesis = each series has unit root and no cointegration. The tests were estimated with the following assumptions: trend assumption (quadratic deterministic trend) and tags interval in first differences. ***1% level of significance, **5% level of significance, *10% level of significance.
Sources: Authors' calculations using the World Inequality Database. 1988–2014. WID.world (accessed December 3, 2018); and World Bank. 1988–2014. World Development Indicators. https://databank.worldbank.org/source/world-development-indicators (accessed December 3, 2018).

Besides using logarithm public expenditure (% of GDP) as the explanatory variable, we included several macroeconomic variables that may influence income inequality. We included logarithm trade (% of GDP), GDP per capita at PPP (current international US dollars), population growth (annual %), oil rents (% of GDP), and tax revenue (% of GDP).

The regression is structured to estimate the following equation:[1]

$$\begin{aligned} inequality_{it} = \alpha_i &+ \beta_1' expense_{it} + \beta_2' \Delta trade_{it} + \beta_3' \Delta gdp_{it} + \beta_4' \Delta pop_{it} \\ &+ \beta_5' \Delta oil_{it} + \beta_6' \Delta tax_{it} + \eta_i t + \theta_t + \varepsilon_{it} \end{aligned} \tag{3}$$

where α_i are individual constants; $\eta_i t$ are individual trends; θ_t is the common time effect; $(1, -\beta_1', -\beta_2', -\beta_3', -\beta_4', -\beta_5', -\beta_6')$ are cointegrating vectors between logarithm public expenditure (% of GDP), trade (% of GDP), GDP per capita at PPP (current international US dollars), population growth (annual %), oil rents (% of GDP), and tax revenue (% of GDP), and ε_{it} is an idiosyncratic error.

One of the key advantages in using the FMOLS and the DOLS estimations is that we can deal with the spurious regression and draw causal effects with the nature of time series that are nonstationary at level. In this case, the standard ordinary least squares (OLS) and the generalized method of moments estimators are inconsistent. McCallum (2010) and Sollis (2011) made a huge contribution in arguing the problems of "spurious regressions." McCallum (2010) suggested so-called spurious regression relationships, which are generally accompanied by clear signs of residual autocorrelation. In our study, the spurious relationships between the series at level $I(1)$, or the nonintegrated variables $inequality_{it}$ and $expense_{it}$, can be resolved by estimating the whole autocorrelation structure. It is

[1]We applied the same identification strategy by using institutional quality ($insti_{it}$) and the interaction term between public expenditure (% of GDP) and institutional quality ($expense \times insti_{it}$) as the explanatory variable in equation (3).

solved in simulations, which result in test statistics closing to true values and not yielding spurious results. However, Sollis (2011) argued that the spurious regression problem can be solved by using an autocorrelation correction. It is shown that if the relevant data generation processes contain higher-order terms, this solution is not as effective as in the first-order case.

In this study, suppose we have two $I(1)$ random vectors with panel observations, $inequality_{it}$ and $expense_{it}$, with large cross section and time series dimensions. By pooling the cross section and time series observations, the strong effect of residuals is attenuated while retaining the signal of $expense_{it}$. In this regard, while the time series is spurious, applying all-time series data in cross sections reduces the limiting variance in a panel regression and provides a consistent estimate of (some) long-run regression coefficient (Malinen 2016). According to Kao and Chiang (2001), the simulations of the sampling behavior show that although the FMOLS estimator provides better estimations than the standard OLS and the generalized method of moments estimators, the DOLS outperforms the other estimations.

The resulting FMOLS estimator is asymptotically unbiased and has the fully efficient mixture normal asymptotics, allowing for standard Wald tests using asymptotic chi-square statistical inference (Sobrado et al. 2014). Complementary to the FMOLS, the panel DOLS is estimated with fixed effects; fixed effects and heterogeneous trends; and fixed effects, heterogeneous trends, and common time effects. The model takes into account cross-sectional dependence by introducing a common time effect, and the estimators are asymptotically normally distributed (Mark and Sul 2003).

In equation (3), although the FMOLS and the DOLS estimators can provide improvements compared to the OLS estimator, we might face other statistical issues—including (i) cointegration between the explanatory variables, (ii) possible endogeneity problem of spurious correlation, and (iii) potential serial correlation—that require us to estimate with great caution.

First of all, the FMOLS and the DOLS estimators do not allow for cointegration between the explanatory variables. In our estimations, we include the leads and lags of the first differences of logarithm trade (% of GDP), GDP per capita at PPP (current international US dollars), population growth (annual %), oil rents (% of GDP), and tax revenue (% of GDP).

Secondly, to address an endogeneity problem of spurious correlation between $inequality_{it}$ and $expense_{it}$, and other explanatory variables, the panel DOLS estimation assumes that μ_{it} is supposed to be correlated most with ρ_i leads and lags of $\Delta expense_{it}$. The possible endogeneity can be controlled by projecting ε_{it} into these p_i leads and lags (Hämäläinen and Malinen 2011):

$$\mu_{it} = \sum_{s=-p_i}^{p_i} \xi'_{i,s} \Delta expense_{i,t-s} + \varepsilon_{it*} = \xi'_i \Delta z_{it} + \varepsilon^*_{it} \tag{4}$$

where z_{it} is a random vector with panel observation, and $\xi_i' z_{it}$ is a vector of projection dimensions. The projection error ε_{it}^* is orthogonal to all leads and lags of $\Delta expense_{it}$, and therefore the estimated equation is transformed as follows:

$$\begin{aligned} inequality_{it} &= \alpha_i + \beta_1' expense_{it} + \beta_2' \Delta trade_{it} + \beta_3' \Delta gdp_{it} + \beta_4' \Delta pop_{it} \\ &\quad + \beta_5' \Delta oil_{it} + \beta_6' \Delta tax_{it} + \eta_i t + \theta_t + \xi_{\mathrm{i}}' z_{it} + \varepsilon_{it} \end{aligned} \tag{5}$$

Finally, to address the potential serial correlation between equilibrium error, ε_{it}, and leads and lags of other cross sections $\Delta expense_{ji}$, $j \neq i$, the panel DOLS computes the same form of second-order asymptotic bias as pooled OLS. Overall, the estimation of equation (5) is consistent under the condition which $T \to \infty$ then $n \to \infty$. Equation (5) therefore can be feasibly estimated in a panel with small to moderate n (Mark and Sul 2003).

We started our regression by estimating individually the impact of public expenditure (% of GDP) ($expense_{it}$), institutional quality ($insti_{it}$), and interaction term between public expenditure (% of GDP) and institutional quality ($expense \times insti_{it}$) on income inequality, measured by the top 1% income share. A few necessary steps were taken: first, we estimated with the constant (level) and no trend; second, we estimated with the constant (level) and trend; finally, we introduced the control variables in our regression, including first differences of logarithm trade (% of GDP) ($\Delta trade_{it}$), GDP per capita at PPP (Δgdp_{it}), annual population growth (Δpop_{it}), oil rents (% of GDP) (Δoil_{it}), and tax revenue (% of GDP) (Δtax_{it}). Only the estimations with the control variables are shown.

Table 5 presents results from the FMOLS and the DOLS estimations using the dataset from 1988 to 2014. For the FMOLS estimation, the long-run variance estimates—Bartlett kernel, Newey–West fixed bandwidth—were used for coefficient covariances. We also used pooled estimation as panel method and fixed leads and lags specification to address the possible endogeneity and serial correlation discussed above. For the DOLS estimation, the same long-run variance estimates were used for coefficient covariances. The pooled estimation as panel method and automatic leads and lags specification were estimated. The first, second, and third leads and lags of the first differences of control variables were estimated as instruments for the explanatory variables. However, only the results from the first leads and lags are shown.

For control variables, trade openness (% of GDP) has positive and statistically significant cointegrating coefficients (significant at 1%) when we estimated with public expenditure (% of GDP) for both FMOLS and DOLS. It becomes negative and statistically significant (significant at 10%) when we estimated with institutional quality and the interaction term between public expenditure (% of GDP) and institutional quality. Per capita GDP shows a mix of direction; yet the majority of coefficients are statistically significant at the 1% or (at least) 5% level. The estimated result of the population growth rate also shows

Table 5. **Public Expenses, Institutional Quality, and Top 1% Income Share**

	Panel Fully Modified Least Squares (FMOLS)			Panel Dynamic Least Squares (DOLS)		
	Model 1	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**
Public expenses (% of GDP)	−0.255**			−0.200*		
	(0.098)			(0.089)		
Institutional quality		−0.102			−1.513*	
		(0.239)			(0.654)	
Public expenses × Institutional quality			−0.776***			−1.043**
			(0.257)			(0.405)
ΔTrade openness (% of GDP)	0.206**	−0.355*	−0.262	0.384***	−0.171	−0.692*
	(0.094)	(0.178)	(0.175)	(0.046)	(0.237)	(0.320)
ΔGDP (per capita at PPP)	−6.024**	9.752***	0.806	−0.637	3.638***	4.223***
	(2.589)	(0.193)	(0.679)	(0.557)	(0.449)	(0.842)
ΔPopulation growth (annual %)	−0.034**	0.070	−0.126	−0.051***	0.143*	0.263*
	(0.016)	(0.220)	(0.085)	(0.009)	(0.063)	(0.112)
ΔOil rents (% of GDP)	0.025	−0.146	−0.072**	−0.087**	−0.010	0.116*
	(0.027)	(0.192)	(0.031)	(0.028)	(0.039)	(0.048)
ΔTax revenue (% of GDP)	0.185*	0.369***	1.485***	−0.327**	0.822***	0.678**
	(0.102)	(0.113)	(0.531)	(0.096)	(0.162)	(0.249)
Adjusted R^2	0.984	0.607	0.890	0.997	0.803	0.693
Countries	8	8	8	8	8	8
Years	1988–2014	1988–2014	1988–2014	1988–2014	1988–2014	1988–2014
Observations	104	49	49	28	18	18

GDP = gross domestic product, PPP = purchasing power parity.

Notes: Data in parentheses indicate standard errors. The regression results were estimated with the FMOLS and the DOLS methods. For the FMOLS, the regressions were estimated with the following assumptions: panel method (pooled estimation); cointegrating equation deterministics (the constant [level] and/or trend); and long-run covariance estimates (Bartlett kernel, Newey–West fixed bandwidth). For the DOLS, the regressions were estimated with the following assumptions: panel method (pooled estimation); cointegrating equation deterministics (the constant [level] and/or trend); and automatic leads and lags specification (based on Schwarz information criterion, max = *). Long-run variance (Bartlett kernel, Newey–West fixed bandwidth) was used for coefficient covariances. ***$p < 0.01$, **$p < 0.05$, and *$p < 0.1$.

Sources: Authors' calculations using the World Inequality Database. 1988–2014. WID.world (accessed December 3, 2018); and World Bank. 1988–2014. World Development Indicators. https://databank.worldbank.org/source/world-development-indicators (accessed December 3, 2018).

a mix of direction. It becomes negative and statistically significant (significant at 1%) when we estimated with public expenditure (% of GDP) for both FMOLS and DOLS; however, it becomes positive and significant at the 5% level when we estimated with institutional quality and the interaction term between public expenditure (% of GDP) and institutional quality for the DOLS. The majority of oil rents (% of GDP) and tax revenue (% of GDP) show only one direction as there is a negative relationship for oil rents (% of GDP) and a positive relationship for tax revenue (% of GDP).

Based on our findings, the globalized forces, as explained by trade (% of GDP), do increase income inequality in the Asia and Pacific countries. Yet, it is not implied that the benefits from trade globalization go only to the rich or the top income earners. It is possible that the living standards of poorer citizens also increase, but not as much as for the rich; therefore, we found the current discontent with globalization in the Asia and Pacific countries is not as intense as in Europe and North America. In the most advanced economies (i.e., European and North American countries), there is a growing belief that globalizing forces are not all good; both ordinary citizens and policy makers think that life was better in the old days and that the fruits of globalization might go only to top earners and the rest of the world (Gray 2017, Willige 2017). According to Shanmugaratnam (2017), various social trends that have occurred in the advanced economies over the last few decades could explain this phenomenon, including stagnant wages, an overall decline in social mobility, a loss of sense of togetherness, and a growing mentality of “us against them.” The estimated results are verified by the Kuznets inverted-U hypothesis. GDP per capita is likely to increase income inequality during the first stage of economic development but decrease it in the long run. According to Blancheton and Chhorn (2019), this result confirms the fact that there is a rising number of people joining the global middle-income class, thanks to an increase in the living standards of people in Asia and the Pacific, especially in India and the PRC, which together account for 36.4% of the global population. The global middle-income class is defined as follows: “[T]hose households with daily expenditures between $10 and $100 per person at PPP. This excludes those who are considered poor in the poorest advanced countries and rich in the richest advanced countries” (Mahbubani 2014, 23). Because rapid demographic growth has enabled strong economic growth, especially in the Asia and Pacific countries, an increase in population has not led to an increase in income inequality. Meanwhile, higher oil rents and higher tax revenue, respectively, decrease and increase income inequality.

Public expenditure (% of GDP) is found to be negative and statistically significant. The estimated value of cointegrating coefficients varies between −0.2552 (significant at 5% for FMOLS) and −0.20004 (significant at 10% for DOLS). Institutional quality is found to be negative and statistically significant with a coefficient of 1.5132 (significant at 10%) only if we estimated with DOLS. When we estimated with the interaction term between public expenditure (% of GDP) and institutional quality, we also found negative and significant coefficients for both FMOLS and DOLS. The results suggest that whenever the Asia and Pacific countries improve their institutional quality enough, higher public expenditures are likely to reduce income inequality.

Compared to countries in Europe and North America, the Asia and Pacific countries have relatively weaker institutional quality. However, as discussed in the previous section, it is likely that better institutional quality does not guarantee a more equal society, or at least weaker institutional quality is not an obstacle to promote the welfare of lower-income citizens. In the modern age of a global single market, even with less effective governance and institutions, some giant or big economies are still able to attain economic growth that is sufficient to allow millions of poor to become middle-income families. For instance, it has been said that the PRC grows because of its government, driven by strong public intervention, while India grows despite its government, driven by market forces even with less effective governance (Mahbubani 2014). Rising trade openness and economic growth in these economies might lead to higher inequality overall, but strong government intervention, through public spending and subsidies, as well as robustly rising incomes help to promote significantly the poor's living standard. In the same way, with impressive progress in higher education and research and development, along with rising social mobility, some authors argue that "*The American Dream Is Alive. In China*" (Hernández and Bui 2018). Considering institutional and political factors in our study, it might be relevant to review the theory of the founding father of economic reform in the PRC, Deng Xiaoping, who said the following: "It doesn't matter whether the cat is black or white, as long as it catches mice" (Li 1977, 107). Therefore, it does not matter whether it is democracy or communism, but whether the political institutions target the majority of the people, especially the poor and the more vulnerable.

B. Granger Causality Tests

Many studies have emphasized that income inequality hurts economic growth, which then leads to greater demand for redistribution through public expenditure and taxes in many societies (Kennedy et al. 2017, Tanninen 1999). This may cause the reverse effect between public expenditure and income inequality. The same logical reasoning is also applied for institutional quality. For example, the interaction of political and income inequality may play a part in blocking the adoption of good institutions (Chong and Gradstein 2007). To address this issue, the Granger causality tests can be statistically applied to estimate whether public expenditure may influence income inequality or vice versa.

In this paper, we used the pairwise Granger causality tests (Granger 1969). We thus estimated the bivariate regressions of the following form:

$$y_t = \alpha_0 + \alpha_1 y_{t-1} + \cdots + \alpha_l y_{t-l} + \beta_1 x_{t-1} + \cdots + \beta_l x_{t-l} + \varepsilon_t \tag{6}$$

$$x_t = \alpha_0 + \alpha_1 x_{t-1} + \cdots + \alpha_l x_{t-l} + \beta_1 y_{t-1} + \cdots + \beta_l y_{t-l} + \varepsilon_t \tag{7}$$

Table 6. **Tests for Granger Noncausality between Public Expenses and Institutional Quality and the Top 1% Income Share**

Explanatory Variable (x)	Dependent Variable (y)	Obs	*F*-Statistic	Prob
Public expenses (% of GDP)	Top 1% income share	150	1.591	0.207
Top 1% income share	Public expenses (% of GDP)		1.339	0.265
Public expenses (% of GDP)	Institutional quality	69	0.649	0.526
Institutional quality	Public expenses (% of GDP)		6.642^{***}	0.002
Countries	8			
Years	1988–2014			

GDP = gross domestic product.
Notes: The null hypothesis is that the explanatory variable (x) does not cause the dependent variable (y). $^{***}p < 0.01$, $^{**}p < 0.05$, and $^{*}p < 0.1$.
Sources: Authors' calculations using the World Inequality Database. 1988–2014. WID.world (accessed December 3, 2018); and World Bank. 1988–2014. World Development Indicators. https://databank.worldbank.org/source/world-development-indicators (accessed December 3, 2018).

where l is a lag length, which corresponds to reasonable beliefs about the longest time over which one of the variables could help predict the other (Granger 1969). From equations (6) and (7), dependent variable y can cause explanatory variable x and, at the same time, explanatory variable x can cause dependent variable y.

The joint null hypotheses of the model are as follow: "y does not Granger-cause x" and "x does not Granger-cause y." We can reject the null hypothesis if the F-statistics, which are the Wald statistics for the joint hypothesis, have a reported p-value at the 1%, 5%, or 10% level of significance.

Table 6 presents the results of Granger noncausality tests between public expenditure (% of GDP) and institutional quality and the top 1% income share in all the Asia and Pacific countries in our dataset. We have no evidence that public expenditure (% of GDP) influences the top 1% income share. It is identical that the influence of the top 1% income share cannot be used to forecast public expenditure as a share of GDP. In the case of institutional quality, we also do not have enough evidence to emphasize that the top 1% income share drives institutional quality; however, we have enough evidence at the 1% level of significance to reject the null hypothesis (i.e., institutional quality would forecast the public expenditure as a share of GDP).

V. Robustness Checks

A. Nonlinearity Analysis

Though government intervention and institutional factors linking to income inequality seem to be linear, the relationship may be generated by different mechanisms at different levels of government intervention and institutional factors. This can lead to thinking about a nonlinear analysis. We included the nonlinear analysis into the methodology, following the studies of Tan and Law (2012), predicting a hump or inverted U-shaped relationship between income inequality and financial factors in line with Shahbaz et al. (2015), studying the Kuznets curve between financial development and income inequality in line with Rojas-Vallejos and Turnovsky (2017), and exploring the nonlinear relationship between tariff reductions and income inequality. The square term of the explanatory variable is included into the equation as follows:[2]

$$\begin{aligned} inequality_{it} &= \alpha_i + \beta_1' expense_{it} + \beta_1^* expense_{it}^2 + \beta_2' \Delta trade_{it} + \beta_3' \Delta gdp_{it} \\ &\quad + \beta_4' \Delta pop_{it} + \beta_5' \Delta oil_{it} + \beta_6' \Delta tax_{it} + \eta_i t + \theta_t + \varepsilon_{it} \end{aligned} \tag{8}$$

From equation (8), the U-shaped nonlinear relationship between public expenditure and inequality predicts $\beta_1' < 0$ and $\beta_1^* > 0$; however, the inverted U-shaped nonlinear relationship predicts $\beta_1' > 0$ and $\beta_1^* < 0$. This is also applied for the institutional quality.

Table 7 shows the estimated results from the FMOLS and DOLS estimations. We followed the same identification strategy as a linear approach in the previous section. We estimated public expenses (% of GDP) and its square value by introducing the control variables in our regression. It is also applied for institutional quality. The long-run variance estimates—Bartlett kernel, Newey–West fixed bandwidth—were used for coefficient covariances for both the FMOLS and DOLS estimations. We also took pooled estimation as panel method and fixed leads and lags specification to address possible endogeneity and serial correlation as discussed above for FMOLS. For the DOLS estimation, the pooled estimation as panel method and automatic leads and lags specification were estimated. The first leads and lags of the first differences of control variables are estimated as instruments for the explanatory variables.

According to the estimated results, public expenditure shows positive and statistically significant cointegrating coefficients at the 1% level for FMOLS and at the 10% level for DOLS. Its square value shows negative and statistically significant cointegrating coefficients at the 1% level for FMOLS and at the 10% level for DOLS. The institutional quality also shows the same direction of coefficient, although the significance level is different. Linking together, we obtained thus the inverted U-shaped nonlinear relationship of public expenditure and institutional quality on income inequality. More precisely, at the early stage of institutional development, a country whose economy has experienced higher public expenditure generates rising income inequality; then, in the long run when the country improves its institutional quality, the higher public expenditure results in lower income inequality.

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[2] We also applied the same identification strategy by using institutional quality ($insti_{it}$) and square institutional quality ($insti_{it}^2$) as explanatory variables.

Table 7. **Nonlinearity Analysis of Public Expenses, Institutional Quality, and the Top 1% Income Share**

	Panel Fully Modified Least Squares (FMOLS)		**Panel Dynamic Least Squares (DOLS)**	
	Model 1	**Model 2**	**Model 3**	**Model 4**
Public expenses (% of GDP)	0.417***		0.109*	
	(0.979)		(0.312)	
Public expenses (% of GDP), squared	−0.117***		−0.319*	
	(0.022)		(0.089)	
Institutional quality		0.882***		0.328
		(0.181)		(1.196)
Institutional quality, squared		−0.322*		−0.251***
		(0.182)		(0.652)
ΔTrade openness (% of GDP)	0.053***	0.049***	0.672**	0.030
	(0.019)	(0.013)	(0.152)	(0.025)
ΔGDP (per capita at PPP)	−0.673***	0.825***	−0.227*	−0.138**
	(0.243)	(0.140)	(0.711)	(0.508)
ΔPopulation growth (annual %)	−1.497	−1.958***	−1.963	0.96***
	(1.598)	(0.733)	(1.987)	(0.298)
ΔOil rents (% of GDP)	−1.670**	0.473	−0.281**	−0.208
	(0.748)	(0.328)	(0.535)	(0.244)
ΔTax revenue (% of GDP)	0.344	0.089	0.870*	0.372***
	(0.251)	(0.138)	(0.235)	(0.040)
Adjusted R^2	0.469	0.964	0.724	0.997
Countries	8	8	8	8
Years	1988–2014	1988–2014	1988–2014	1988–2014
Observations	106	92	16	35

GDP = gross domestic product, PPP = purchasing power parity.
Notes: Data in parentheses indicate standard errors. The regression results were estimated with the FMOLS and the DOLS methods. For the FMOLS, the regressions were estimated with the following assumptions: panel method (pooled estimation); cointegrating equation deterministics (the constant [level] and/or trend); and long-run covariance estimates (Bartlett kernel, Newey–West fixed bandwidth). For the DOLS, the regressions were estimated with the following assumptions: panel method (pooled estimation); cointegrating equation deterministics (the constant [level] and/or trend); and automatic leads and lags specification (based on Schwarz information criterion, max = *). Long-run variance (Bartlett kernel, Newey–West fixed bandwidth) was used for coefficient covariances. $^{***}p < 0.01$, $^{**}p < 0.05$, and $^{*}p < 0.1$.
Sources: Authors' calculations using the World Inequality Database. 1988–2014. WID.world (accessed December 3, 2018); and World Bank. 1988–2014. World Development Indicators. https://databank.worldbank.org/source/world-development-indicators (accessed December 3, 2018).

B. Alternative of Measuring Inequality

Although this study brings new insight into the thinking of inequality using a new measurement of the top income segment, it may face bias in that the top 1% income share cannot capture the effect of public spending and institutional quality in promoting the economic opportunity of the poor and the middle-income class. To see the complete picture, we also used the SWIID (version 8.2) as the robustness check (Solt 2019). Notice that the SWIID is the Gini index of inequality in the equalized household market (pretax, pretransfer income).[3]

[3]For details, see Fredrick Solt. "Using the SWIID in Stata." https://osf.io/tj7ck/download.

We applied the same identification strategy for both the linear and nonlinear long-run approaches for the estimations of the top 1% income share. In Table 8, we estimated separately public expenditure and institutional quality, the squares of public expenditure and institutional quality, and the interaction term between public expenditure and institutional quality. We obtained higher values for adjusted R-squared and the number of observations. This is likely due to having a more complete SWIID database compared to the top 1% income share. The SWIID dataset is available for nearly all of our eight sample countries in Asia and the Pacific.

We obtained nearly similar results as estimating with the top 1% income share, considering the direction and significance level of the cointegrating coefficients. We presumed therefore that public expenditure and institutional quality drive inequality reduction, and that the effects follow the inverted U-shaped nonlinear relationship in the long run.

VI. Conclusion

Inequality has indeed mattered not only in the past but also in the present and the future. Therefore, the legitimacy of this issue has always been in the equation. Many studies have linked inequality to government intervention and institutional quality, but most of them were not quantitatively estimated to understand the long-run equilibrium relationship. Thus, the main objective of our paper is to examine the significance of such a long-run relationship, in both linear and nonlinear analysis, by applying the strength of FMOLS and DOLS, as well as the Granger causality tests. We used a dataset for eight countries in Asia and the Pacific—Australia, India, Japan, Malaysia, the PRC, the Republic of Korea, Singapore, and Thailand—from 1998 to 2014.

As reported by our estimated results, there are negative long-run, steady-state effects of government intervention (measured by public expenditure as a share of GDP) and institutional quality (measured by the WGI) on income inequality (measured by the top 1% income share in the World Inequality Database first developed by Piketty and Zucman [2014]) in the sample countries in Asia and the Pacific. The effect of institutional quality has only a one-way Granger causality link to income inequality. The existence of a nonlinear relationship between public expenditure and institutional factors linking to income inequality is also found. It implies that, at the early stage of institutional development, a country whose economy has experienced higher public expenditure generates rising income inequality; then, in the long run when a country improves its institutional quality, the higher public expenditure results in lower income inequality. The findings also suggested a nonlinear relationship in the long run when we estimated results with the Gini index of inequality of the SWIID (version 8.2).

To develop a full picture of how government intervention and institutional factors influence inequality in the long run, additional studies are needed. Firstly,

Table 8. **Public expenses (% of GDP), Institutional Quality, and Inequality by SWIID version 8.2**

	Panel Fully Modified Least Squares (FMOLS)					Panel Dynamic Least Squares (DOLS)				
	Model 1	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**	**Model 7**	**Model 8**	**Model 9**	**Model 10**
Public expenses (% of GDP)	−0.087*** (0.027)			−1.508*** (0.502)		−0.064 (0.945)			0.327* (0.167)	
Public expenses (% of GDP), squared				0.032** (0.013)					−0.087* (0.049)	
Institutional quality		−0.414*** (0.056)			0.649*** (0.153)		−0.046** (0.018)			0.512*** (0.143)
Institutional quality, squared					−0.276* (0.156)					−1.214 (1.736)
Public expenses × Institutional quality			−0.412*** (0.056)					−0.044** (0.0197)		
ΔTrade openness (% of GDP)	0.116*** (0.015)	−0.096* (0.049)	−0.107** (0.052)	0.026** (0.011)	0.040*** (0.010)	0.150 (0.102)	0.304*** (0.022)	0.297*** (0.026)	0.047*** (0.012)	0.026* (0.013)
ΔGDP (per capita at PPP)	0.276** (0.105)	2.037*** (0.237)	2.616*** (0.272)	0.670*** (1.577)	3.711** (1.446)	0.696 (0.810)	1.107*** (0.066)	1.192*** (0.086)	0.920*** (0.272)	3.022* (1.548)
ΔPopulation growth (annual %)	0.024*** (0.006)	−0.038 (0.033)	−0.006 (0.036)	0.945 (0.733)	0.141 (0.693)	0.056 (0.056)	−0.013 (0.020)	−0.005 (0.022)	1.874 (1.566)	0.303 (0.760)
ΔOil rents (% of GDP)	−0.036*** (0.007)	0.019** (0.010)	0.013 (0.011)	−0.590** (0.272)	−0.245 (0.216)	−0.0266 (0.160)	−0.041*** (0.012)	−0.039*** (0.013)	−0.524* (0.305)	−0.516 (0.350)
ΔTax revenue (% of GDP)	0.033 (0.033)	0.312*** (0.113)	0.466*** (0.131)	0.034 (0.138)	0.135 (0.118)	−0.006 (0.693)	0.218*** (0.035)	0.229*** (0.041)	0.003 (0.114)	0.0240 (0.102)
Adjusted R–squared	0.984	0.499	0.485	0.960	0.977	0.989	0.990	0.990	0.993	0.987
Countries	8	8	8	8	8	8	8	8	8	8
Years	1988–2014	1988–2014	1988–2014	1988–2014	1988–2014	1988–2014	1988–2014	1988–2014	1988–2014	1988–2014
Observations	121	65	65	127	107	65	36	36	60	88

GDP = gross domestic product; PPP = purchasing power parity.

Notes: Data in parentheses indicate standard errors. The regression results were estimated with the FMOLS and the DOLS methods. For the FMOLS, the regressions were estimated with the following assumptions: panel method (pooled estimation); cointegrating equation deterministics (the constant [level] and/or trend); and long-run covariance estimates (Bartlett kernel, Newey–West fixed bandwidth). For the DOLS, the regressions were estimated with the following assumptions: panel method (pooled estimation); cointegrating equation deterministics (the constant [level] and/or trend); and automatic leads and lags specification (based on Schwarz information criterion, max = *). Long-run variance (Bartlett kernel, Newey–West fixed bandwidth) was used for coefficient covariances. $^{***}p < 0.01$, $^{**}p < 0.05$, and $^{*}p < 0.1$.

Sources: Authors' calculations using the World Inequality Database. 1988–2014. WID.world (accessed December 3, 2018); and World Bank. 1988–2014. World Development Indicators. https://databank.worldbank.org/source/world-development-indicators (accessed December 3, 2018).

it might be possible to use other tools of public intervention through government expenditure at a more disaggregated level, which are extensively studied in short- and medium-run analyses. Secondly, we could compare the Asia and Pacific countries to other countries like those in Latin America that have had similar economic and political development paths. Finally, while using the average values of the WGI, we have not taken a closer look at their six subcategories because each dimension can be subject to a different explanation of inequality. While the average score of the WGI is higher, it does not mean that these subcategories are all equally higher. It should thus be subject to further investigation as institutional quality at the very first level of aggregation might not be rational enough to differentiate its effect.

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Appendix

Table A.1. **Control Variable Definitions and Sources**

Variable	Description and ID	Sources
Trade (% of GDP)	Trade is the sum of exports and imports of goods and services measured as a share of gross domestic product. ID: NE.TRD.GNFS.ZS	World Bank national accounts data, OECD national accounts data files
GDP per capita at PPP (current international US dollars)	GDP per capita based on PPP is GDP converted to international dollars using PPP rates. An international dollar has the same purchasing power over GDP as the US dollar has in the United States. GDP at purchaser’s prices is the sum of gross value added by all resident producers in the economy plus any product taxes and minus any subsidies not included in the value of the products. It is calculated without making deductions for depreciation of fabricated assets or for depletion and degradation of natural resources. Data are in current international dollars based on the 2011 ICP round. ID: NY.GDP.PCAP.PP.CD	World Bank ICP database
Population growth (annual %)	Annual population growth rate for year t is the exponential rate of growth of midyear population from year $t-1$ to t, expressed as a percentage. Population is based on the de facto definition of population, which counts all residents regardless of legal status or citizenship. ID: SP.POP.GROW	Derived from total population. Population source: (1) United Nations Population Division. World Population Prospects: 2019 Revision, (2) Census reports and other statistical publications from national statistical offices, (3) Eurostat: Demographic Statistics, (4) United Nations Statistical Division. Population and Vital Statistics Report (various years), (5) US Census Bureau: International Database, and (6) Secretariat of the Pacific Community: Statistics and Demography Programme.

Continued.

Table A.1. *Continued.*

Variable	Description and ID	Sources
Oil rents (% of GDP)	Oil rents are the difference between the value of crude oil production at world prices and total costs of production. ID: NY.GDP.PETR.RT.ZS	Estimates based on sources and methods described in World Bank. 2011. *The Changing Wealth of Nations: Measuring Sustainable Development in the New Millennium*. Washington, DC.
Tax revenue (% of GDP)	Tax revenue refers to compulsory transfers to the central government for public purposes. Certain compulsory transfers such as fines, penalties, and most social security contributions are excluded. Refunds and corrections of erroneously collected tax revenue are treated as negative revenue. ID: GC.TAX.TOTL.GD.ZS	International Monetary Fund, Government Finance Statistics Yearbook and data files, and World Bank and OECD GDP estimates.

GDP = gross domestic product, ICP = International Comparison Program, PPP = purchasing power parity, OECD = Organisation for Economic Co-operation and Development, US = United States.
Sources: Authors' calculations using the World Inequality Database. 1988–2014. WID.world (accessed December 3, 2018); and World Bank. 1988–2014. World Development Indicators. https://databank.worldbank.org/source/world-development-indicators (accessed December 3, 2018).

Table A.2. **Pairwise Correlation among Control Variables**

	trade	gdp	pop	oil	tax
Trade (% of GDP) [trade]	1				
Significance level					
Observation	216				
Per capita GDP (at PPP) [gdp]	0.4*	1			
Significance level	0.0				
Observation	200	200			
Population growth (annual %) [pop]	0.4	−0.02	1		
Significance level	0.5	0.8			
Observation	216	200	216		
Oil rents (% of GDP) [oil]	0.3	−0.1	0.4*	1	
Significance level	0.3	0.1	0.0		
Observation	173	161	173	173	
Tax revenue (% of GDP) [tax]	0.1	0.4*	0.2*	0.1	1
Significance level	0.4	0.0	0.01	0.3	
Observation	191	179	191	148	191

GDP = gross domestic product, PPP = purchasing power parity.
Note: *1% level of significance.
Sources: Authors' calculations using the World Inequality Database. 1988–2014. WID.world (accessed December 3, 2018); and World Bank. 1988–2014. World Development Indicators. https://databank.worldbank.org/source/world-development-indicators (accessed December 3, 2018).

Table A.3. **Variance Inflation Factor among Control Variables**

Variable	VIF	1/VIF
Public expenses (% of GDP)	5.7	0.2
Tax revenue (% of GDP)	4.6	0.2
Per capita GDP (at PPP)	4.4	0.2
Population growth (annual %)	3.95	0.3
Oil rents (% of GDP)	3.2	0.3
Trade (% of GDP)	2.4	0.4
Mean VIF	4.04	

GDP = gross domestic product, PPP = purchasing power parity, VIF = variance inflation factor.
Note: Top 1% income share is used as dependent variable.
Sources: Authors' calculations using the World Inequality Database. 1988–2014. WID.world (accessed December 3, 2018); and World Bank. 1988–2014. World Development Indicators. https://databank.worldbank.org/source/world-development-indicators (accessed December 3, 2018).

Table A.4. **List of Countries**

Country	Region	Income Status	Institutional Status
Australia	East Asia and Pacific	High income	Strong quality
People's Republic of China	East Asia and Pacific	Upper-middle income	Weak quality
India	South Asia	Lower-middle income	Weak quality
Japan	East Asia and Pacific	High income	Strong quality
Republic of Korea	East Asia and Pacific	High income	Strong quality
Malaysia	East Asia and Pacific	Upper-middle income	Strong quality
Singapore	East Asia and Pacific	High income	Strong quality
Thailand	East Asia and Pacific	Upper-middle income	Weak quality

Source: Authors' compilation.

Table A.5. **Eight Countries in Asia and the Pacific**

	Obs	Mean	Median	Max	Min	Std Dev	Skewness	Kurtosis
Top 1% income share	178	12.0	10.6	23.5	0.0	4.6	1.0	3.3
Public expenses (% of GDP)	205	17.0	15.8	26.8	10.8	4.1	1.0	3.1
Institutional quality	152	0.6	0.5	1.7	−0.6	0.8	−0.0	1.5
Trade openness (% of GDP)	216	106.7	54.9	439.7	13.3	109.1	1.6	4.3
GDP per capita	200	4.1	4.2	4.9	3.0	0.5	−0.6	2.6
Population growth	216	1.3	1.2	5.3	−1.5	0.9	0.8	4.5
Oil rents (% of GDP)	173	1.5	0.9	9.6	0.0	1.9	1.9	2.0
Tax revenue (% of GDP)	191	14.2	13.7	24.9	8.1	4.2	0.9	3.1

GDP = gross domestic product.
Notes: Asia and the Pacific comprises Australia, India, Japan, Malaysia, the People's Republic of China, the Republic of Korea, Singapore, and Thailand. Data are from 1988 to 2014.
Sources: Authors' calculations using the World Inequality Database. 1988–2014. WID.world (accessed December 3, 2018); and World Bank. 1988–2014. World Development Indicators. https://databank.worldbank.org/source/world-development-indicators (accessed December 3, 2018).