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# FINANCIAL DEVELOPMENT AND THE FDI-INEQUALITY NEXUS

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*This article empirically investigates whether the link between foreign direct investment (FDI) and income inequality varies with financial development. Using a smooth transition regression model to a panel of developing and advanced countries over the period of 1976–2005, the results indicate that financial development indeed defines the relationship between FDI and inequality. FDI raises income inequality and the effect becomes stronger in magnitude with financial sophistication. The results also indicate a large variation in the FDI effect across countries and over time, contingent on financial development.* (*JEL* C23, F40, O15)

## I. INTRODUCTION

Rising income inequality coinciding with increasing financial globalization, especially in foreign direct investment (FDI), in most developed countries and some developing economies over the past decades has led researchers to explore the link between FDI and income inequality. This shift from the growth effect to the inequality impact of FDI is of particular importance as it helps devise policy measures that can allow the efficiency and growth gains of FDI to be shared more equally across all segments of a society.

The question of whether FDI affects income inequality is subject to considerable dispute. Nevertheless, little consensus has been reached. Theoretically, such a link is difficult to build as there are plausible models suggesting equalizing effects from FDI liberalization (e.g., Mundell 1957) as well as models suggesting disequalizing effects (e.g., Bornschier and Chase-Dunn 1985; Feenstra and Hanson 1996; Gaston and Nelson 2002).

Empirically, problems often arise either from lack of available and comparable data on inequality in international panels or because of endogeneity and reverse causality. In the empirical literature, the Deininger and Squire (1996) database is the standard reference for inequality studies; however, the coverage of this database is sparse and unbalanced, and consequently its measures of inequality originate from different sources and refer to a variety of income and population definitions.[1] For instance, many cross-country studies on inequality have used the Deininger-Squire-based World Income Inequality Database (WIID). The Gini coefficients in WIID are based on different income definitions (income/expenditure; gross/net), different recipient units (individuals/households), and population coverage (urban/rural/all). Even when adjustments are made to improve data comparability, these differences may still result in serious data inconsistency. This poses important problems of comparability, which may undermine the robustness of the results.[2] On

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1. Deininger and Squire (1996) collected many disparate surveys of income and expenditure inequality and compiled them into a single panel, offering 693 country/year observations since 1947.

2. Atkinson and Brandolini (2001) present a critique of the Deininger and Squire database that focuses, in part, on the fact that many different types of data drawn from different sources are mixed up in the data set. In general, they criticize the use of secondary statistics and show how both cross-country comparisons and time-series analyses may crucially depend on the choice of data.

**ABBREVIATIONS**

FDI: Foreign Direct Investment
GDP: Gross Domestic Product
M&A: Merger and Acquisition
OLS: Ordinary Least Squares
PSTR: Panel Smooth Transition Regression
PTR: Panel Threshold Regression
SWIID: Standardized World Income Inequality Database
WIID: World Income Inequality Database



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the other hand, the decision to liberalize FDI and redistribute income/wealth may be political (Acemoglu, Johnson, and Robinson 2005; Acemoglu and Robinson 2006; Quinn 2000). As such, correlation between FDI and inequality may be driven by the common factor or simply reflect the preferences of policy-makers or political incumbents.

This article empirically investigates the link between FDI and income inequality. Our analysis centers on the FDI-inequality relationship at different stages of financial development. Instead of focusing on understanding why FDI is detrimental or beneficial to income distribution, we focus on understanding the second-order relationship. Given that FDI may have an effect on income inequality, we intend to assess the implications of this relationship when financial arrangements are continually changing. Thus, our contribution lies in recognizing that the role of FDI and inequality in society evolves with the financial system.

This investigation is critical and has important policy implications. Given a growing literature finding the growth and efficiency benefits of FDI at the higher levels of financial development, if inequality-reducing effect of FDI is found at the latter stages of financial development, it implies that a sufficient level of financial development is necessary for everyone in the society to benefit from FDI to the same extent. This calls for financial sector reform policy to materialize the benefits from FDI. In contrast, if financial development strengthens the inequality-enhancing effect of FDI, it indicates that financial development allows only the rich and the politically connected to exploit the new growth opportunities created by FDI globalization. Such exclusive growth may not be sustainable as rising inequality can lead to a backlash against FDI liberalization and protectionist pressures.

The possibility that the effect of FDI on inequality depends on the level of financial development clearly corresponds to the definition of a threshold regression model. We thus address this issue employing a panel smooth transition regression (PSTR) model with fixed effects introduced by González, Teräsvirta, and van Dijk (2005) and Fouquau, Hurlin, and Rabaud (2008). The PSTR model is a regime-switching model that allows a smooth transition between regimes depending on the value of the threshold (transition) variable (here, financial development). The basic idea underlying this model is that when some threshold level of financial development is passed, the economy smoothly moves to another regime where the FDI-inequality relationship is qualitatively different from that of the previous regime. Because the transition variable is individual-specific and time-varying, the regression coefficients for each of the individuals in the panel are changing over time. The PSTR methodology hence provides a simple parametric approach that allows for the capture of both the cross-country heterogeneity and time instability of the impact of FDI on income inequality contingent on the level of financial development in the panel framework.[3]

Using Gini coefficients of household net income in the Standardized World Income Inequality Database (SWIID) recently developed by Solt (2009) as our preferred inequality measure, we can construct a (balanced) panel from 1976 through 2005 with more observations on within-country income inequality than other studies in this area do. We find that the relationship between FDI and income inequality is governed by a two-regime model with income inequality increasing with FDI in both regimes but at an increasing rate in the second regime with higher financial development. The evidence implies that FDI raises inequality in income distribution, which is further strengthened by financial development.

The remainder of the article is organized as follows. Section II gives a brief review of the literature on the FDI-inequality link and describes the potential role played by financial development in the link. Section III introduces the smooth transition regression model and describes the data. Section IV reports the empirical results, and Section V concludes the analysis.

## II. A BRIEF LITERATURE REVIEW

### *A. FDI and Income Inequality*

The existing research does not unanimously establish the consequences of FDI for income inequality.[4] Mundell (1957) argues that multinational corporations bring capital into a developing country with scarce capital, decreasing the relative return on capital

3. The PSTR approach hence generalizes the Hansen (1999) threshold regression model, assuming there is a discrete jump in the inequality-FDI link in terms of financial development. It is also different from the conventional interaction methodology, assuming FDI to be a linear function of financial development in their interaction term.

4. It is noted that since most studies in the literature focus on the inequality effect of inward FDI, this section reviews theoretical and empirical works of this line.

to labor. Thus foreign capital competes with domestic capital for domestic labor, increasing labor income and decreasing the profitability of domestic firms. This effect would speed up convergence of income of labor relative to capital, thereby decreasing (increasing) income inequality in capital-scarce developing countries (capital-abundant developed countries). By contrast, others predict that FDI strengthens income inequality across countries when considering imperfect competition, outsourcing activity, and technology heterogeneity. As claimed in Hymer (1976), foreign firms have advantages over domestic firms in access to proprietary knowledge, management skills, and economies of scale, and these foreign firms choose to locate in a country to turn this advantage into profits, rather than exporting their products or licensing domestic firms to produce for them. And multinational corporations often use more advanced technologies, employ more skilled workers, and pay a wage premium over local firms, producing highly paid elites and large number of marginalized workers. Therefore, countries that are wholly dependent on foreign capital would experience increasing income inequality (Bornschier and Chase-Dunn 1985). Likewise, in Feenstra and Hanson (1996) and Gaston and Nelson (2002), FDI increases the relative demand for skilled workers in both developed and developing countries as the North outsources relatively unskilled intensive products to the South, where these are relatively skilled intensive.

The inequality increasing effect may be further strengthened if FDI accelerates skill-bias technological change and hence increases the demand for skilled workers. The skill-bias technology may spill over from subsidiaries of multinationals to domestic firms and may take place through demonstration and/or imitation (domestic firms imitate new technologies of foreign firms), competition (entrance of foreign firms leads pressure on domestic firms to adjust their activities and to introduce new technologies), linkages (spillovers through transactions between multinationals and domestic firms), and/or training (domestic firms upgrade the skill of their employees to enable them to work with new technologies).

Consistent with the division in the theoretical literature on the distributional repercussion of FDI, empirical analyses reach diverse conclusions. In a cross-country context, Tsai (1995) studies the link between FDI and income inequality using a sample of 33 developing countries and finds that FDI increases inequality only in some Asian countries. However, Alderson and Nielsen (1999) find that controlling for the different geographical regions does not change the significant positive effect of foreign capital penetration on income inequality. Gopinath and Chen (2003) find, with a sample of 11 developing countries, that FDI flows widen the skilled–unskilled wage gap for a subset of developing countries although they appear to lead to cross-country convergence of wages. Basu and Guariglia (2007) use a panel of around 80 countries to test a theoretical model linking FDI to growth and inequality in human capital and conclude that inward FDI promotes economic inequality. Others such as Mahler, Jesuit, and Roscoe (1999), Sylwester (2005), and Adam (2008) find that FDI variables are not statistically significant in explaining income inequalities.

In country-specific studies, while Zhang and Zhang (2003) find that increasing FDI inflows contribute to greater income inequality for China, Wei, Yao, and Liu (2009) blame the uneven distribution of FDI (rather than FDI itself) to be the cause of rising regional inequality in China. Jensen and Rosas (2007) show a decrease in income inequality of Mexico with FDI inflows whereas Mah (2003) finds no significant effect for Korea. Blonigen and Slaughter (2001) fail to find any significant effects of FDI on wage inequality between skilled and unskilled workers for the United States. In contrast, Chintrakarn, Herzer, and Nunnenkamp (2012) find that FDI exerts a significant and robust negative effect on income inequality in the United States, but with much heterogeneity across states.

The inconclusive empirical evidence may be suggestive of nonlinearity in the link between FDI and inequality. Taylor and Driffield (2005) find that inward FDI increases wage inequality but at a decreasing rate over time in the United Kingdom. Figini and Görg (2011) find that for developing countries wage inequality increases with FDI, but this effect diminishes with further increases in FDI. In contrast, wage inequality decreases with FDI for developed countries. Lin, Kim, and Wu (2013) find that FDI increases inequality in income distribution when a country achieves a threshold of human capital between 6.0 and 6.7 years of secondary schooling; below this threshold, however, FDI improves income distribution. In this article, we consider financial development as a potential vehicle in shaping the nonlinearity in the link between FDI and inequality.

*B. The Role of Financial Development*

Financial development may be relevant because financial development has nonlinear or nonmonotonic effect on income inequality, constitutes a source of comparative advantage, and/or materializes international spillovers of technology and knowledge embodied in FDI. Firstly, the importance of well-functioning financial systems in income inequality has been established in the literature. In the presence of investment indivisibilities and locally increasing returns to scale, financial imperfections arising from informational asymmetries and transaction costs produce credit constraints that are particularly binding for small enterprises and the poor that lack collateral, credit histories, and connections. By easing financial frictions and therefore by allowing more poor people and entrepreneurs to access and obtain external finance, financial development improves the allocation of capital and alleviates income inequality (Aghion and Bolton 1997; Banerjee and Newman 1993; Galor and Zeira 1993). Financial development can operate not only on the extensive marginal but also on an intensive marginal. According to the new political economy literature (Acemoglu, Johnson, and Robinson 2005; Rajan and Zingales 2003b), it is primarily the rich and politically connected who benefit from improvements in the financial system. This is not only because they have sufficient wealth for collateral (dubbed the tyranny of collateral), but also because the rich are able to prevent small firms from accessing external finance and reduce the ability of the poor to improve their economic well-being. Thus, financial development deepens, rather than broadens, the access to credit, thereby widening the gap between the poor and the rich.

Still, some argue that the effect of financial development may be nonlinear or nonmonotonic. For example, in Greenwood and Jovanovic (1990), financial intermediaries arise endogenously to mitigate informational asymmetries. Because the organization of financial intermediaries is costly at early stages of development, only the rich can access and benefit from better financial markets. However, at the latter stages of development, financial structure becomes more extensive and income inequality across the rich and the poor declines because financial development helps an increasing proportion of the society. Kim and Lin (2011) provide empirical evidence of this sort.

Second, while differences in financial development across countries are one of the significant determinants of FDI flows, FDI can facilitate development of financial systems. The two-way link between FDI and financial development thus opens up another possible channel through which financial development has nonlinear effects on the FDI-inequality relationship. On the one hand, the extent of financial development determines investment flows. Di Giovanni (2005) shows how domestic financial deepening affects firms investing abroad. Using a panel dataset of cross-border merger and acquisition (M&A) deals for 1990–1999, he finds that deep financial markets in the acquisition countries can play a significant role in cross-border M&As. Klein, Peek, and Rosengren (2002) argue that the collapse of the banking sector in Japan played a significant role in reducing the amount of FDI from Japan to the United States in the 1990s, even after controlling for the relative wealth movements caused by fluctuations in stock prices and exchange rates. Campos and Kinoshita (2010) find a stronger effect on FDI from financial sector reforms than from privatization and trade liberalization, suggesting that foreign investors do value highly a host country's financial system that is able to allocate capital efficiently, monitor firms, ameliorate, diversify and share risk, and ultimately mobilize savings.

On the other hand, financial openness leads to development of financial systems through several channels. Financial openness enhances the functioning of domestic financial systems by intensifying competition and importing financial services (Klein and Olivei 2008; Levine 2001). Financial liberalization increases the efficiency level of the financial system by weeding out inefficient financial institutions and creating greater pressure on the domestic government to institute reforms (Claessens, Demirgüç-Kunt, and Huizinga 2001; Stiglitz 2000; Stulz 1999). Financial integration also affects financial development by allowing access to foreign financial markets in the form of direct lending by foreign intermediaries and listing on foreign stock markets (Masten, Fabrizio, and Masten 2008). And finally, opening to trade and capital flows weakens the relative political power and the incentives of incumbent industrialized firms or financial intermediaries to block financial development in order to reduce entry and competition (Rajan and Zingales 2003a).

Third, the literature on FDI and economic growth emphasizes a country's absorptive capacity in reaping gains from FDI. In addition to sheer foreign capital it supplies, FDI encourages the adoption of new technology in the production process through technological spillovers, and stimulates knowledge transfers in terms of labor training and skill acquisitions, introduction of alternative management practice, and better organizational arrangements. However, such spillovers are not automatic because local conditions have important effects in influencing firms' adoption and implementation of foreign technologies and skills. Particularly, it is argued that the economy's ability to take advantage of potential FDI spillovers depends upon the development of local financial markets. For example, in Alfaro et al. (2004), the successful acquisition of new technologies introduced by foreign firms will generally involve a process of reorganization and reinvestment by their domestic competitors. To the extent that this process is externally financed from domestic sources, efficient financial markets will enhance the competitive response of the domestic industry. Evidently, Hermes and Lensink (2003), Alfaro et al. (2004, 2010), Durham (2004), Prasad, Rajan, and Subramanian (2007), Alfaro, Chanda, and Sayek (2009), and Campos and Kinoshita (2010) show that countries with a well-developed financial market gain significantly from FDI.[5]

In summary, neither existing theoretical studies nor empirical ones have rigorously examined the effects of improvement in the financial sector on the FDI-inequality relationship. Our analysis does not take a position on whether financial development affects FDI and/or income inequality independently, but focuses more narrowly on the question of whether financial development itself affects the marginal relationship between FDI and inequality, after controlling for other factors.

## III. MODEL SPECIFICATION AND DATA

### A. *Model Specification*

To evaluate the contribution of FDI to income inequality across countries and over time, we estimate the following dynamic regression equation[6]:

$$\text{(1)}\quad \text{ineq}_{it} = \alpha_i + \beta_{it}\ \text{fdi}_{it-1} + \delta'_{it} w_{it-1} + \epsilon_{it}$$

where ineq is an income inequality indicator, fdi is an FDI indicator, $w$ is a set of control variables, and $\alpha_i$ is a country fixed effect. $i = 1, 2, \ldots, N$ is the country indicator, $t = 1, 2, \ldots, T$ is the time period index, and $\epsilon$ is the error term.

In a panel context, it is typically assumed that $\beta_{it} = \beta$ and $\delta_{it}' = \delta'$ for all $i = 1, \ldots, N$ and $t = 1, \ldots, T$. Such a poolability assumption is somewhat restrictive as there are substantial differences between countries in their institutions, macroeconomic policies, redistributive schemes, and economic conditions, which may cause differences in the sensitivity of inequality to FDI across countries. Moreover, Equation (1) implies that the effect of FDI on income inequality is constant over the time period considered in the model. This assumption appears to be misleading especially when examining large time dimension panels as well as when the economy experiences structural shifts due to internal or external shocks.

The PSTR methodology can solve heterogeneity and time variability problems simultaneously by introducing threshold effects in a linear panel model specification. Following González, Teräsvirta, and van Dijk (2005) and Fouquau, Hurlin, and Rabaud (2008), the two-regime PSTR model takes the form:

(2)
$$\begin{aligned}\text{ineq}_{it} &= \alpha_i + \beta_0 \text{fdi}_{it-1} + \beta_1 \text{fdi}_{it-1} g\left(q_{it-1}; \gamma, c\right)\\ &\quad + \delta_0 w_{it-1} + \delta_1 w_{it-1} g\left(q_{it-1}; \gamma, c\right) + \epsilon_{it}\end{aligned}$$

where $g(q_{it-1}; \gamma, c)$ is a transition function of the threshold variable $q_{it-1}$, which is financial development, and is continuous and bounded between 0 and 1.

As in González, Teräsvirta, and van Dijk (2005) and Fouquau, Hurlin, and Rabaud (2008), we consider the following transition function of logistic specification:

(3)
$$g\left(q_{it-1}; \gamma, c\right) = \left\{1 + \exp\left[-\gamma \prod_{j=1}^{m} (q_{it-1} - c_j)\right]\right\}^{-1},$$
$$\gamma > 0,\ c_1 < c_2 < \ldots < c_m$$

5. Kim, Lin, and Suen (2013) find that FDI benefits countries with less developed financial systems, in terms of domestic investment, however.

6. There is a potential issue of endogeneity in the inequality-FDI link. However, Fouquau, Hurlin, and Rabaud (2008) show that the PSTR limits the potential endogeneity bias, because, for each level of threshold variable, there is a particular value of the estimated regression parameter. Despite so, to mitigate the endogeneity problem, we use one-period (3 years) lag of FDI, financial development, and other controls.

where $c = (c_0, c_1, \ldots, c_m)'$ denotes a $m$-dimensional vector of location parameters and $q_{it-1}$ indicates the threshold level at which the transition function reaches an inflexion point. Parameter $\gamma$ determines the slope of the transition function, that is, the speed of the transition from one regime to another. Furthermore, the restrictions $\gamma > 0$ and $c_1 < c_2 < \ldots < c_m$ are imposed for identification. As common in the literature on threshold panel data analysis, this study assumes $m = 1$ or $m = 2$ to capture the nonlinearities due to regime switching. Note that the case $m = 1$ refers to a logistic PSTR model, and $m = 2$ corresponds to a logistic quadratic PSTR specification. When $\gamma \to \infty$, the transition function $g(q_{it-1}; \gamma, c)$ tends to be an indicator function, that is, $g(q_{it-1}; \gamma, c) = 0$ if $q_{it-1} < c$, and $g(q_{it-1}; \gamma, c) = 1$ if $q_{it-1} \geq c$. Furthermore, the transition is sharp as in the panel threshold regression (PTR) model developed by Hansen (1999). When $\gamma \to 0$, the transition function $g(q_{it-1}; \gamma, c)$ becomes constant and the model collapses into a homogenous or linear panel regression model with fixed effects (a so-called "within" model).

The influence of FDI on income inequality is then defined as a weighted average of parameters $\beta_0$ and $\beta_1$. For a given threshold variable $q_{it-1}$, the effect of FDI on income inequality for the $i$th country at time $t$ is:

(4)

$$\frac{\partial \text{ineq}_{it}}{\partial \text{fdi}_{it-1}} = \beta_0 + \beta_1 g\left(q_{it-1}; \gamma,\ c\right),$$

$$\text{with} \begin{cases} \beta_0 + \beta_1 \leq \frac{\partial \text{ineq}_{it}}{\partial \text{fdi}_{it-1}} \leq \beta_0 & \textit{if } \beta_1 < 0 \\ \beta_0 \leq \frac{\partial \text{ineq}_{it}}{\partial \text{fdi}_{it-1}} \leq \beta_0 + \beta_1 & \textit{if } \beta_1 > 0 \end{cases}.$$

Note that parameters $\beta_0$ and $\beta_1$ do not correspond to the effect of FDI on income inequality. Parameter $\beta_0$ corresponds to the FDI coefficient only if the transition function $g(q_{it-1}; \gamma, c)$ tends toward 0. The sum of the $\beta_0$ and $\beta_1$ parameters corresponds to the FDI coefficient only if the transition function $g(q_{it-1}; \gamma, c)$ tends toward 1. Between these two extremes, the FDI coefficient is defined as a weighted average of parameters $\beta_0$ and $\beta_1$. Therefore, it is important to note that it is generally difficult to directly interpret the values of these parameters (as in a probit or logit model). It is generally preferable to interpret (1) the sign of these parameters, which indicates an increase or a decrease in the FDI coefficient depending on the value of the threshold variable and (2) the varying coefficient in the time and individual dimensions given by Equation (4).

Note that the PSTR model can be generalized to $r + 1$ extreme regimes as follows:

(5)

$$\begin{aligned} \text{ineq}_{it} &= \alpha_i + \beta_0 \text{fdi}_{it-1} \\ &\quad + \sum_{j=1}^{r} \beta_j \text{fdi}_{t-1} g\left(q_{_{it-1}}; \gamma_j,\ c_j\right) + \delta_0 w_{it-1} \\ &\quad + \sum_{j=1}^{r} \delta_j w_{it-1} g\left(q_{_{it-1}}; \gamma_j,\ c_j\right) + \epsilon_{it} \end{aligned}$$

where $r$ is the number of transition functions. The marginal impact of FDI on income inequality is then given by:

$$\frac{\partial ineq_{it}}{\partial fdi_{it-1}} = \beta_0 + \sum_{j=1}^{r} \beta_j g\left(q_{it-1}; \gamma_j,\ c_j\right). \tag{6}$$

The estimation of the parameters of the PSTR model consists of eliminating the country fixed effects $\alpha_i$ by removing country-specific means and then estimating the transformed model using nonlinear least squares (for details, see González, Teräsvirta, and van Dijk 2005). Following Colletaz and Hurlin (2006) and Fouquau, Hurlin, and Rabaud (2008), we adopt a three-step procedure for estimating the PSTR model. First, we test for linearity against the PSTR model. Then, if linearity is rejected, we determine the number of transition functions. Finally, we remove country-specific means using the standard fixed-effects transformation and then apply nonlinear least squares to estimate the parameters of the transformed model.

Testing for linearity in Equation (2) can be performed by testing $H_0 : r = 0$, which expresses no regime-switching effect in our data. However, if the PSTR model under the null hypothesis is unidentified, classical tests will not follow a standard distribution, the so-called Davies Problem (Davies 1977).

To solve this problem, the transition function can be replaced by its first-order Taylor expansion around $\gamma = 0$, and an equivalent hypothesis should be tested in an auxiliary regression. The auxiliary regression based on $m = 1$ and $m = 2$, respectively, can be written as:

$$\begin{aligned} \text{ineq}_{it} &= \alpha_i + \theta_0 z_{it-1} + \theta_1 z_{it-1} q_{it-1} + \epsilon^*_{it} \\ \text{ineq}_{it} &= \alpha_i + \theta_0 z_{it-1} + \theta_1 z_{it-1} q_{it-1} \\ &\quad + \theta_2 z_{it-1} q^2_{it-1} + \epsilon^*_{it} \end{aligned} \tag{7}$$

where $z = (\text{fdi}, w)'$ .$\epsilon^*_{it} = \epsilon_{it} + R\left(q_{it-1}; \gamma, c\right)$ and $R(q_{it-1}; \gamma, c)$ are the remainder of the Taylor expansion. Testing linearity (no regime-switching effect) against the PSTR model means simply testing the null hypothesis $H_0 : \theta_1 = \theta_2 = 0$ in this auxiliary regression, which is a linear model. If we denote $\text{SSR}_0$ as the panel sum of squared residuals under $H_0$ (linear panel model with individual effects) and $\text{SSR}_1$ as the panel sum of squared residuals under $H_1$ (PSTR model with two regimes), the corresponding $F$-statistic is given by:

$$(8) \quad \text{LM}_\text{F} = \frac{\left[\left(\text{SSR}_0 - \text{SSR}_1\right)/mK\right]}{\left[\text{SSR}_0/\left(\text{TN} - N - m(K+1)\right)\right]}$$

where $K$ is the number of explanatory variables. Under the null hypothesis, the $\text{LM}_\text{F}$ has an asymptotic $F(mK,\ \text{TN} - N - m(K+1))$ distribution.

Similar methodology is used regarding testing the number of transition functions in the model, or, equivalently, the number of extreme regimes. If linearity is rejected, a sequential approach is used to test the null hypothesis of no remaining nonlinearity in the transition function. Testing for no remaining nonlinearity consists of checking whether there is one transition function ($H_0 : r = 1$) or at least two transition functions ($H_1 : r = 2$) defined as:

$$(9) \quad \begin{aligned} \text{ineq}_{it} &= \alpha_i + \beta_0 \text{fdi}_{it-1} + \delta_0 w_{it-1} \\ &\quad + \left(\beta_1 \text{fdi}_{it-1} + \delta_1 w_{it-1}\right) g\left(q_{it-1}; \gamma_1,\ c_1\right) \\ &\quad + \left(\beta_1 \text{fdi}_{it-1} + \delta_1 w_{it-1}\right) g\left(q_{it-1}; \gamma_2,\ c_2\right) \\ &\quad + \epsilon_{it}. \end{aligned}$$

As in the case of one transition function, the test consists of replacing the second transition function by its first-order Taylor expansion around $\gamma_2 = 0$ and then testing linear constraints on the parameters. Using this first-order Taylor expansion, the model becomes:

$$(10) \quad \begin{aligned} \text{ineq}_{it} &= \alpha_i + \theta_0 z_{it-1} + \theta_1 z_{it-1} g\left(q_{it-1}; \gamma_1, c_1\right) \\ &\quad + \theta_2 z_{it-1} q_{it-1} + \epsilon^*_{it} \\ \text{ineq}_{it} &= \alpha_i + \theta_0 z_{it-1} + \theta_1 z_{it-1} g\left(q_{it-1}; \gamma_1, c_1\right) \\ &\quad + \theta_2 z_{it-1} q_{it-1} + \theta_3 z_{it-1} q^2_{it-1} + \epsilon^*_{it}. \end{aligned}$$

Let $\text{SSR}_0$ denote the panel sum of squared residuals under $H_0$ (the PSTR model with one transition function) and $\text{SSR}_1$ the sum of squared residuals under $H_1$ (the PSTR model with at least two transition functions). The test statistics can be calculated in the same manner as above. The sequential procedure is then as follows. Given a PSTR model with $r = r^*$, test the null $H_0 : r = r^*$ against $H_1 : r = r^* + 1$. If $H_0$ is not rejected, the procedure stops. Otherwise, the null hypothesis $H_0 : r = r^* + 1$ is tested against $H_1 : r = r^* + 2$. The testing procedure continues until the first acceptance of the null hypothesis of no remaining nonlinearity occurs.

In addition, linearity tests also serve to determine the appropriate order of $m$ in the logistic transition function (Equation (3)). Teräsvirta (1994) proposed a sequence of tests for choosing between $m = 1$ and $m = 2$. Within the PSTR framework, this testing sequence runs as follows. Using the auxiliary regression for the case of $r = 1$ in Equation (7), test the null hypothesis $H_0 : \theta_2 = \theta_1 = \theta_0 = 0$. If it is rejected, test $H^3_0 : \theta_2 = 0$. Then, exclude $\theta_2 = 0$ and test $H^2_0 : \theta_1 = 0 \left|\theta_2 = 0\right.$ and $H^2_0 : \theta_0 = 0 \left|\theta_1 = \theta_2 = 0\right.$. These hypotheses are tested by the ordinary $F$ tests, denoted by $F_3, F_2$, and $F_1$, respectively. The decision rule is as follows: the $m = 2$ transition function is selected in cases where the $p$ value corresponding to $F_2$ is the smallest, and the $m = 1$ transition function is chosen for other cases.

We then eliminate the country fixed effects $\alpha_i$ by removing country-specific means. The country means in Equation (2) are expressed as follows:

$$(11) \quad \begin{aligned} \overline{\text{ineq}_i} &= \alpha_i + \beta_0 \overline{\text{fdi}_i} + \beta_1 \overline{x_i}(\gamma, c) + \delta_0 \overline{w_i} \\ &\quad + \delta_1 \overline{y_i}(\gamma, c) + \overline{\epsilon_i} \end{aligned}$$

where $\overline{\text{ineq}_i}$, $\overline{\text{fdi}_i}$, $\overline{x_i}(\gamma,\ c)$, $\overline{w_i}$, $\overline{y_i}(\gamma,\ c)$, and $\overline{\epsilon_i}$ are country means. Subtracting Equation (11) from Equation (2) yields:

$$(12) \quad \widetilde{\text{ineq}}_{it} = \beta' \widetilde{\text{fdi}}_{it-1}(\gamma, c) + \delta' \widetilde{w}_{it-1}(\gamma,\ c) + \widetilde{\epsilon}_{it}$$

where $\widetilde{\text{ineq}}_{it} = \text{ineq}_{it} - \overline{\text{ineq}_i}$, $\widetilde{\text{fdi}}_{it}(\gamma, c) = \left(\widetilde{\text{fdi}}_{it} - \overline{\text{fdi}_i}, \text{fdi}_{it} g\left(q_{it-1}; \gamma, c\right) - \overline{x_i}(\gamma, c)\right)$, $\widetilde{w}_{it}(\gamma, c) = \left(\widetilde{w}_{it} - \overline{w_i}, w_{it} g\left(q_{it-1}; \gamma, c\right) - \overline{y_i}(\gamma, c)\right)$, $\beta' = \left(\beta_0, \beta_1\right)$, $\delta' = \left(\delta_0, \delta_1\right)$, and $\widetilde{\epsilon}_{it} = \epsilon_{it} - \overline{\epsilon}_i$.

Let $\phi = (\beta', \delta')$ and $\widetilde{S}_{it}(\gamma, c) = \left(\widetilde{\text{fdi}}_{it\text{-}1}(\gamma, c), \widetilde{w}_{it\text{-}1}(\gamma, c)\right)$. Then Equation (12) can be written as:

$$(13) \quad \widetilde{\text{ineq}}_{it} = \phi' \widetilde{S}_{it}(\gamma, c) + \widetilde{\epsilon}_{it}.$$

Then, given the pair $(\gamma, c)$, the estimate is obtained using ordinary least squares (OLS) as follows:

$$\widehat{\phi}(\gamma, c) = \left[\sum_{i=1}^{N}\sum_{t=1}^{T}\widetilde{S}_{it}(\gamma, c)\widetilde{S}_{it}(\gamma, c)'\right]^{-1} \times \left[\sum_{i=1}^{N}\sum_{t=1}^{T}\widetilde{S}_{it}(\gamma, c)\,\mathrm{in\widetilde{e}q}_{it}\right]$$

Conditional on $\widehat{\phi}(\gamma, c)$, the parameters of the transition function $c$ and $\gamma$ are estimated by minimizing the concentrated sum of squared errors via nonlinear least squares:

$$(14)\quad \left(\widehat{\gamma}, \widehat{c}\right) = \mathrm{ArgMin}\sum_{i=1}^{N}\sum_{t=1}^{T} \times [\mathrm{in\widetilde{e}q}_{it} - \widehat{\phi}'(\gamma, c)\,\widetilde{S}_{it}(\gamma, c)]^2.$$

## *B. Data*

Our sample consists of 42 developing and developed countries. All of the data, which span the period 1976–2005, are derived from different sources and averaged over ten nonoverlapping 3-year periods: 1976–1978, 1979–1981, …, 2000–2002, and 2003–2005.[7] The selection of countries is restricted by the availability of data on inequality. Limited overlap between data for FDI and financial development, and the requirement, imposed by the PSTR methodology, of balanced data reduce the sample size drastically.

Our inequality measure, the Gini coefficient, is obtained from a global inequality dataset—the Solt (2009) SWIID database—which has been made available very recently and which ensures data comparability both through time and across countries. The SWIID standardizes the WIID, with the Luxembourg Income Study as a standard, while minimizing reliance on problematic assumptions by using as much information as possible from proximate years within the same country. Specifically, SWIID improves data availability and comparability for cross-national research by exploiting the fact that different types of Gini coefficients display systematic relationships. The Gini coefficient for gross income is typically larger than the coefficient for net income, which in turn is larger than the Gini coefficient for expenditure. Similarly, Gini coefficients for households are typically lower than coefficients for individuals. The difference between gross and net income Gini coefficients depends on the degree to which taxes and transfers are progressive and redistribute income from rich to poor, and hence should vary across countries and within countries over time. Instead of adding three points to net income-based inequality observations to make them comparable with the gross income-based observations of Deininger and Squire, or including dummy variables to correct for different types of Gini coefficients in the same regression, Solt (2009) estimates the ratios between different types of Gini coefficients, relying on information about the ratio in the same country close in time to increase the number of comparable observations. The SWIID currently incorporates comparable Gini indices of gross and net income inequality for 173 countries for as many years as possible since 1960. It also includes estimates of the uncertainty of these statistics. We use the *net income Gini coefficient* (*gini_net*) from Solt (2009) as our preferred distributional measure and dependent variable. As a robustness check, we also use the *gross income Gini coefficient* (*gini_gross*), as a dependent variable.

Concerning the indicators of FDI openness, because FDI liberalization includes not only inward FDI but also outward FDI, we first use inward FDI as our measure of the extent of FDI liberalization because most theoretical arguments and empirical investigations focus on inward FDI. We also experiment with the sum of inward and outward FDI, as a share of GDP, as a robustness check. These measures are from Lane and Milesi-Ferretti's (2007) External Wealth of Nations Database and are expressed in natural logarithms.

Regarding the threshold variable, because there is no single aggregate index of financial development in the literature, we use principal component analysis to produce a new aggregate index. Ideally, the principal component analysis should be based on indicators from the banking sector, stock markets, and bond markets so as to capture different aspects of financial development. However, the availability of data on stock

7. To allow for varying coefficients over time within cross sections might be problematic when there seems to be memory in the data, that is, the series show properties of nonstationarity. FDI-to-GDP ratios are cases in point. Averaging the data over 3 years and reducing the sample to $t = 10$ does not solve the problem of unit roots in the data since the data generating process remains unchanged. There is the risk of performing spurious regressions in the presence of nonstationary variables. We then perform panel unit root tests. The evidence shows that all variables considered are stationary. This mitigates the aforementioned concern. We are grateful to one anonymous referee for pointing out this.

market and bond market development is severely limited before 1975 or even later, so the analysis focuses on financial intermediary development. The measure is based on three widely used indicators of financial intermediary development and denoted as *findev*. (1) Liquid Liabilities, calculated as the liquid liabilities of banks and nonbank financial intermediaries (currency plus demand and interest-bearing liabilities) over GDP. It measures the size, relative to the economy, of financial intermediaries including three types of financial institutions: the central bank, deposit money banks, and other financial institutions. (2) Private Credit, defined as the credit issued to the private sector by banks and other financial intermediaries divided by GDP, excluding the credit issued to government, government agencies, and public enterprises, as well as the credit issued by the monetary authority and development banks. This captures general financial intermediary activities provided to the private sector. (3) Commercial-Central Bank, the ratio of commercial bank assets over the sum of commercial bank and central bank assets. It proxies the advantages of financial intermediaries in channeling savings to investment, monitoring firms, exerting corporate governance, and undertaking risk management relative to the central bank. Since these indicators are used to measure the size of the banking system, *findev* mainly captures the size of bank-based intermediation. *findev* is the first principal component of these three indicators described above and accounts for 70% of their variation. The weights resulting from principal component analysis over the period 1976–2005 are 0.56 for Liquid Liabilities, 0.67 for Private Credit, and 0.52 for Commercial-Central Bank. We also experiment with private credit as a sensitivity test. The data on these indicators are obtained from the World Bank, Financial Structure and Economic Development Database (2007).

To strengthen our empirical results, we include additional control variables. These variables are the (logarithm of) initial value of income inequality for each period as inequality tends to change slowly over time, the (logarithm of) real per-capita GDP growth (*rgdppc_gr*) to account for the impact of economic growth on distribution, and the (logarithm of) average years of secondary school attainment (*schooling*) as a proxy for the effect of human capital on inequality. Included also are the (logarithm of the) value of credit advanced by financial intermediaries to the private sector divided by GDP (*findev*) to proxy for the effect of financial development on inequality, the (logarithm of the) ratio of government spending to GDP (*gov_spending*) to account for the provision of public goods, the degree of intervention in the marketplace, and the possible use of redistributive expenditures. Finally, the (logarithm of) inflation (*inflation*), measured by percentage changes in the consumer price index, is included to proxy for the effects of macroeconomic instability on inequality, and the (logarithm of the) level of trade share, measured as the sum of exports and imports as a percentage of GDP (*trade*), is added to account for the effect of trade openness on inequality. All controls, except for the initial inequality measures, are obtained from the World Development Indicators (2012) of the World Bank and are measured at the initial values in each period to ameliorate endogeneity. Table A1 of the Appendix provides information on the variables used in the analysis.

## IV. EMPIRICAL RESULTS

### A. *Main Results*

To test for the existence of thresholds in the FDI-inequality link, financial development is measured by the aggregate index of financial intermediary development (*findev*) at the beginning of each period and treated as a threshold variable. Table 1 reports the results of the linearity test and the specification test of no remaining nonlinearity. Three different model specifications are considered for two alternative FDI indicators: inward FDI and total FDI. Model 1 includes initial inequality and initial real per-capita GDP growth for each period as controls; Model 2 adds initial inflation, initial trade share, initial government spending, and initial human capital for each period; and Model 3 considers initial financial development per period.[8] For each model, we compute $LM_F$ statistics for the linearity tests ($H_0$: $r = 0$ vs. $H_1$: $r = 1$) and for the tests of no remaining nonlinearity ($H_0$: $r = a$ vs. $H_1$: $r = a + 1$). The values of the statistics are reported until the first acceptance of $H_0$. As shown, across alternative model specifications and different FDI indicators, the null hypothesis of linearity can be rejected at 0.1 significance level or lower, suggesting that the relationship between FDI and

8. As explained by Fouquau, Hurlin, and Rabaud (2008), the threshold variable may have a direct effect on the dependant variable. In this case, one could misleadingly find switching. To examine this possibility we conduct a test of no remaining linearity with direct effects, with the threshold variable used as an explanatory variable, in Model 3.

**TABLE 1**
$LM_F$ Tests for Linearity and Remaining Nonlinearity

Dependent variable: *gini_net*
Threshold variable: *findev*

| | m = 1 | | | m = 2 | | |
|---|---|---|---|---|---|---|
| | **Model 1** | **Model 2** | **Model 3** | **Model 1** | **Model 2** | **Model 3** |
| Panel A: Inward FDI | | | | | | |
| $H_0: r=0$ vs. | 2.136* | 1.758* | 2.047** | 2.257** | 1.864** | 1.939** |
| $H_1: r=1$ | (0.095) | (0.095) | (0.040) | (0.037) | (0.029) | (0.016) |
| $H_0: r=1$ vs. | 1.876 | 1.365 | 1.479 | 2.476 | 0.725 | 0.856 |
| $H_1: r=2$ | (0.133) | (0.219) | (0.163) | (0.023) | (0.749) | (0.620) |
| Observations (N*T) | 420 | 420 | 420 | 420 | 420 | 420 |
| Panel B: Total FDI | | | | | | |
| $H_0: r=0$ vs. | 4.550*** | 1.841* | 1.778* | 2.738** | 3.180*** | 1.887** |
| $H_1: r=1$ | (0.004) | (0.078) | (0.080) | (0.013) | (0.000) | (0.020) |
| $H_0: r=1$ vs. | 0.452 | 1.096 | 1.089 | 1.654 | 0.495 | 0.751 |
| $H_1: r=2$ | (0.716) | (0.365) | (0.370) | (0.131) | (0.935) | (0.741) |
| Observations (N*T) | 420 | 420 | 420 | 420 | 420 | 420 |

*Notes*: The testing procedure works as follows. First, test a linear model ($r=0$) against a model with one threshold ($r=1$). If the null hypothesis is rejected, test the single threshold model against a double threshold model ($r=2$). The procedure is continued until the hypothesis no additional threshold is not rejected. The $LM_f$ statistics of linearity tests are reported for alternative model specifications. The corresponding *p* values are reported in parentheses. Control variables in Model 1 include initial inequality and initial real per-capita GDP growth for each period; Model 2 adds initial inflation, initial trade share, initial government spending, and initial human capital for each period; and Model 3 considers initial financial development.

***Significant at 1%; **significant at 5%; *significant at 10%.

**TABLE 2**
Tests for Choosing the PSTR Model for the Period 1976–2005

Dependent variable: *gini_net*
Threshold variable: *findev*

| | Inward FDI | | | Total FDI | | |
|---|---|---|---|---|---|---|
| **$r^*=1$** | **Model 1** | **Model 2** | **Model 3** | **Model 1** | **Model 2** | **Model 3** |
| $H_0^3: \theta_2 = 0$ | 0.88 | 1.69 | 1.22 | 1.97 | 2.40** | 1.35 |
| $F_3$ | (0.4517) | (0.1097) | (0.2867) | (0.1179) | (0.0208) | (0.2197) |
| $H_0^2: \theta_1 = 0 \mid \theta_2 = 0$ | 0.08 | 1.24 | 1.35 | 0.38 | 1.06 | 1.60 |
| $F_2$ | (0.9698) | (0.2794) | (0.2161) | (0.7650) | (0.3880) | (0.1241) |
| $H_0^1: \theta_0 = 0 \mid \theta_1 = \theta_2 = 0$ | 1127.42*** | 495.50*** | 432.75*** | 1304.85*** | 571.50*** | 499.76*** |
| $F_1$ | (0.0000) | (0.0000) | (0.0000) | (0.0000) | (0.0000) | (0.0000) |
| Observations (N*T) | 420 | 420 | 420 | 420 | 420 | 420 |

*Notes*: $F_3$, $F_2$, and $F_1$ have an asymptotic F($mK$, $TN-N-m(K+1)$) distribution. The PSTR model with $m=2$ is chosen if the rejection of $H_0^2$ is the strongest one, otherwise the PSTR model with $m=1$ is chosen. *p* values are in parentheses.

***Significant at 1%; **significant at 5%; *significant at 10%.

income inequality is nonlinear and depends on the level of financial development. Thus, using a linear panel model in which income inequality is assumed to be homogenous across countries over time may lead to fallacious estimates, as the estimated coefficient could vary from one country to another and change over time in response to potential structural changes in the economy.

Additionally, as Table 1 indicates, the null hypothesis of two extreme regimes ($r=1$) cannot be rejected at the conventional level except for the case of inward FDI in Model 1 with $m=2$. This outcome suggests that in a PSTR model, a small number of extreme regimes is sufficient to capture nonlinearity, that is, cross-country heterogeneity and its time variability in the FDI-inequality relationship. Recall that a smooth transition model, even with two extreme regimes ($r=1$), can be viewed as a model with an infinite number of intermediate regimes. The FDI coefficients are defined in each period and for each country as weighted averages of the values obtained in the two extreme regimes. The weights depend on the value of the transition function. Thus, even if $r=1$, this model would allow for a continuum of coefficient values (or

**TABLE 3**
Parameter Estimates for the Final PSTR Models

Dependent variable: *gini_net*
Threshold variable: *findev*

| | Inward FDI | | | Total FDI | | |
|---|---|---|---|---|---|---|
| | **Model 1** | **Model 2** | **Model 3** | **Model 1** | **Model 2** | **Model 3** |
| Initial FDI: $\beta_0$ | 0.0105** | 0.0079** | 0.0070** | 0.0092*** | 0.0062*** | 0.0036* |
| | (0.0033) | (0.0034) | (0.0036) | (0.0020) | (0.0022) | (0.0022) |
| $\beta_1$ | 0.0062 | 0.0140** | 0.0130* | 0.0340*** | 0.0146** | 0.0128* |
| | (0.0075) | (0.0073) | (0.0076) | (0.0098) | (0.0075) | (0.0077) |
| *Initial gini:* $\delta_0$ | 0.8384*** | 0.8260*** | 0.8250*** | 0.8187*** | 0.8402*** | 0.8608*** |
| | (0.0311) | (0.0311) | (0.0316) | (0.0300) | (0.0272) | (0.0264) |
| $\delta_1$ | −0.0084 | −0.0176 | −0.0069 | −0.0512*** | −0.0356 | −0.0450 |
| | (0.0094) | (0.0209) | (0.0240) | (0.0141) | (0.0229) | (0.0299) |
| *Rgdppc_gr:* $\delta_0$ | 0.0007 | −0.0007* | −0.0007 | −0.0012*** | −0.0009** | −0.0006 |
| | (0.0008) | (0.0004) | (0.0004) | (0.0004) | (0.0004) | (0.0004) |
| $\delta_1$ | −0.0045** | 0.0025** | 0.0019 | 0.0062 | 0.0025 | 0.0005 |
| | (0.0021) | (0.0014) | (0.0013) | (0.0043) | (0.0016) | (0.0014) |
| *Inflation:* $\delta_0$ | | −0.0057 | 0.0005 | | −0.0029 | 0.0094 |
| | | (0.0095) | (0.0106) | | (0.0093) | (0.0119) |
| $\delta_1$ | | 0.0300 | −0.0658 | | −0.0335 | −0.1274 |
| | | (0.1104) | (0.1060) | | (0.0841) | (0.0894) |
| *Trade:* $\delta_0$ | | −0.0094 | −0.0095 | | −0.0083 | −0.0079 |
| | | (0.0080) | (0.0080) | | (0.0083) | (0.0085) |
| $\delta_1$ | | −0.0440*** | −0.0425*** | | −0.0353** | −0.0205 |
| | | (0.0126) | (0.0123) | | (0.0159) | (0.0133) |
| *Gov_spending:* $\delta_0$ | | −0.0105 | −0.0106 | | −0.0044 | −0.0043 |
| | | (0.0112) | (0.0112) | | (0.0065) | (0.0066) |
| $\delta_1$ | | 0.0478*** | 0.0466** | | 0.0965*** | 0.0581*** |
| | | (0.0178) | (0.0187) | | (0.0357) | (0.0227) |
| *Schooling:* $\delta_0$ | | 0.0023 | 0.0008 | | 0.0158 | 0.0113 |
| | | (0.0133) | (0.0136) | | (0.0136) | (0.0127) |
| $\delta_1$ | | 0.0294** | 0.0405 | | −0.0298 | 0.0119 |
| | | (0.0244) | (0.0282) | | (0.0356) | (0.0235) |
| *Findev:* $\delta_0$ | | | 0.0058 | | | 0.0108 |
| | | | (0.0069) | | | (0.0074) |
| $\delta_1$ | | | −0.0204 | | | −0.0035 |
| | | | (0.0186) | | | (0.0183) |
| Location parameters c | 2.7481 | 2.7702 | 2.7542 | 2.8346 | 2.7945 | 2.7306 |
| Slopes parameters | 4.0394 | 8.3805 | 6.8649 | 22.1534 | 6.5373 | 4.8156 |
| Sum of squared residuals | 0.2535 | 0.2230 | 0.2224 | 0.3036 | 0.2122 | 0.1893 |
| AIC | −7.3528 | −7.4234 | −7.4117 | −7.1726 | −7.4729 | −7.5727 |
| BIC | −7.2759 | −7.2695 | −7.2385 | −7.0957 | −7.3190 | −7.3995 |
| Observations (N*T) | 420 | 420 | 420 | 420 | 420 | 420 |

*Note*: The standard errors in parentheses are corrected for heteroskedasticity.
***Significant at 1%; **significant at 5%; *significant at 10%.

regimes), each associated with a different value of the transition function $g(.)$ between 0 and 1.

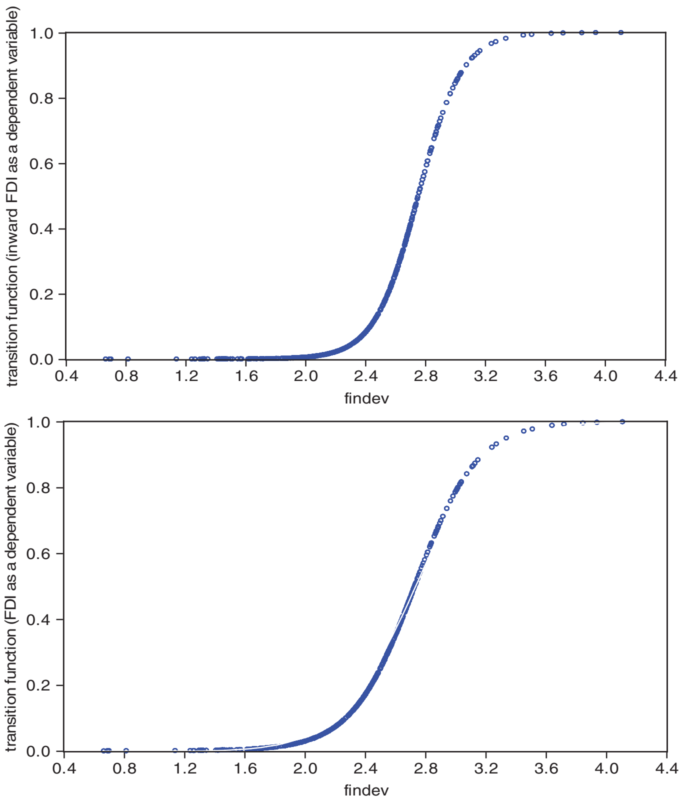
We then examine the PSTR model with $m=1$ or $m=2$. Table 2 reports the results. As illustrated, across different models and FDI indicators, $F_1$, rather than $F_2$, is the largest statistic, meaning that the PSTR model with $m=1$ better captures the nonlinearity in the relationship between FDI and income inequality.[9]

Table 3 reports the parameter estimates of the final PSTR models with $r=1$ and $m=1$. All reported PSTR models lead to positive estimates of the adjustment speed $\gamma$, and the estimates of the location parameter $c$, the turning point of the transition function, are well within the bounds of the transition variable. The estimated slope parameter is small and ranges from 4.8156 to 22.1534, implying that the transition from one regime to another is continuous and smooth. The location parameter $c$ ranges from 2.7306 to 2.8346, which divides the sample into two broad regimes.

9. The PSTR model is a generalization of a two-regime threshold model, while the PSTR model with $m=1$ is a generalization of a single-regime-switching model (Haug and Siklos 2006). In a two-regime threshold model, the adjustment to deviations from the threshold value takes place once the deviation reaches the upper or lower threshold.

**FIGURE 1**

Estimated Transition Function of the PSTR Model against Financial Development for Model 3 of Table 3.




*Note:* $y$ axis is the transition function $g(q_{it};\ \gamma,\ c)$ and the $x$ axis is the transition variable: aggregate index of financial intermediary development (*findev*).

When the transition variable (i.e., *findev*) takes on values less than the estimated threshold values ($q_{it-1} < c$), the transition function approaches to zero ($g(q_{it-1}; \gamma, c) \to 0$) and hence the coefficient of FDI is given by $\beta_0$. We call this regime a low-financial-development regime. When the transition variable exceeds the estimated threshold values ($q_{it-1} \geq c$), however, the transition function approaches to one ($g(q_{it-1}; \gamma, c) \to 1$) and hence the coefficient of FDI is given by $\beta_0 + \beta_1$. We call this regime a high-financial-development regime. As illustrated, across alternative model specifications and different FDI indicators, the coefficient estimates $\widehat{\beta}_0$ and $\widehat{\beta}_1$ are positive and statistically significant. The impacts of FDI on inequality ($\widehat{\beta}_0$ and $\widehat{\beta}_0 + \widehat{\beta}_1$) are significantly positive in two regimes, suggesting that as FDI increases, income inequality slowly rises at early stages of financial development; this pattern smoothly changes around the financial development level of 2.7306–2.8346, and thereafter income inequality rises rapidly.

Figure 1 shows the shape of the estimated logistic function versus the transition variable (*findev*) based on Model 3 in Table 3 for inward FDI and total FDI, respectively. Regimes with low-financial-development country time observations (for which $g(.) = 0$) and high-financial-development observations (for which $g(.) = 1$) are identified, together with a transition phase from one regime to the other. As shown, the change between regimes is quite gradual. These are indicated by the estimated transition parameters $\widehat{\gamma} = 6.8649$ and 4.8156, respectively, for inward FDI and total FDI cases. The estimated threshold values of 2.7542 and 2.7306 point to the half way of the transition, meaning that when $q_{it\text{-}1} = c$, $g(q_{it-1;}\ \gamma,\ c) = 1/2$. It indicates the half-way point between the

**TABLE 4**
Parameter Estimates for the Final PSTR Models

Dependent variable: *gini_net*
Threshold variable: *private credit*

| | Inward FDI | | | Total FDI | | |
|---|---|---|---|---|---|---|
| | **Model 1** | **Model 2** | **Model 3** | **Model 1** | **Model 2** | **Model 3** |
| Panel A: Tests for linearity and remaining nonlinearity | | | | | | |
| $m=1$ | | | | | | |
| $H_0: r=0$ vs. $H_1: r=1$ | 1.620 | 2.004** | 3.292*** | 2.114* | 1.905* | 1.770* |
| | [0.140] | [0.017] | [0.000] | [0.098] | [0.068] | [0.082] |
| $H_0: r=1$ vs. $H_1: r=2$ | 1.439 | 1.274 | 0.798 | 2.044 | 1.711 | 1.661 |
| | [0.199] | [0.221] | [0.688] | [0.107] | [0.105] | [0.107] |
| $m=2$ | | | | | | |
| $H_0: r=0$ vs. $H_1: r=1$ | 0.221* | 2.366*** | 3.137*** | 3.688** | 2.009* | 1.833* |
| | [0.058] | [0.004] | [0.000] | [0.012] | [0.053] | [0.070] |
| $H_0: r=1$ vs. $H_1: r=2$ | 2.213** | 1.808** | 0.562 | 1.474 | 1.595 | 1.666 |
| | [0.041] | [0.036] | [0.911] | [0.221] | [0.135] | [0.105] |
| Panel B: Tests for choosing the PSTR model | | | | | | |
| $H_0^3: \theta_2=0$ | 0.35 | 3.16*** | 4.81*** | 0.86 | 3.08*** | 4.50*** |
| $F_3$ | [0.7904] | [0.0030] | [0.0000] | [0.4627] | [0.0037] | [0.0000] |
| $H_0^2: \theta_1=0 \mid \theta_2=0$ | 2.30* | 1.81* | 2.14** | 1.94 | 1.56 | 1.72* |
| $F_2$ | [0.0771] | [0.0838] | [0.0318] | [0.1224] | [0.1463] | [0.0913] |
| $H_0^1: \theta_0=0 \mid \theta_1=\theta_2=0$ | 817.29*** | 357.23*** | 314.03*** | 1059.08*** | 460.21*** | 402.82*** |
| $F_1$ | [0.0000] | [0.0000] | [0.0000] | [0.0000] | [0.0000] | [0.0000] |
| Panel C: Parameter estimates | | | | | | |
| Initial FDI: $\beta_0$ | 0.0089** | 0.0085** | 0.0087** | 0.0041** | 0.0090*** | 0.0038* |
| | (0.0039) | (0.0036) | (0.0035) | (0.0018) | (0.0032) | (0.0022) |
| $\beta_1$ | −0.0013 | 0.0114** | 0.0105* | 0.0101** | 0.0148** | 0.0161** |
| | (0.0040) | (0.0060) | (0.0064) | (0.0044) | (0.0067) | (0.0068) |
| Location parameters c | 3.3917 | 4.5475 | 4.5790 | 4.4422 | 4.5761 | 4.5773 |
| Slopes parameters $\gamma$ | 9.3979 | 7.0451 | 4.8817 | 8.5726 | 13.3632 | 7.7725 |
| Sum of squared residuals | 0.2543 | 0.2413 | 0.2402 | 0.2064 | 0.2145 | 0.1922 |
| AIC | −7.3496 | −7.3443 | −7.3347 | −7.5584 | −7.4621 | −7.5577 |
| BIC | −7.2726 | −7.1904 | −7.1615 | −7.4815 | −7.3082 | −7.3845 |
| Observations (N*T) | 420 | 420 | 420 | 420 | 420 | 420 |

*Note*: The standard errors ($p$ values) in parentheses (brackets) are corrected for heteroskedasticity.
***Significant at 1%; **significant at 5%; *significant at 10%.

low- and high-financial-development regimes for 42 countries. It is worth noting that the transition phase includes a substantial number of observations, while there are only a few observations in regime 2 when $g(.)=1$.[10] This outcome is most likely due to the large degree of heterogeneity among countries.[11]

10. We also check whether our results are driven by outliers (countries with too high or low average levels of financial development). First, we drop Japan and Malaysia (countries with the highest average level of financial development) and then two more countries, Hungary and Sri Lanka (countries with the lowest average level of financial development). We find qualitatively similar results.

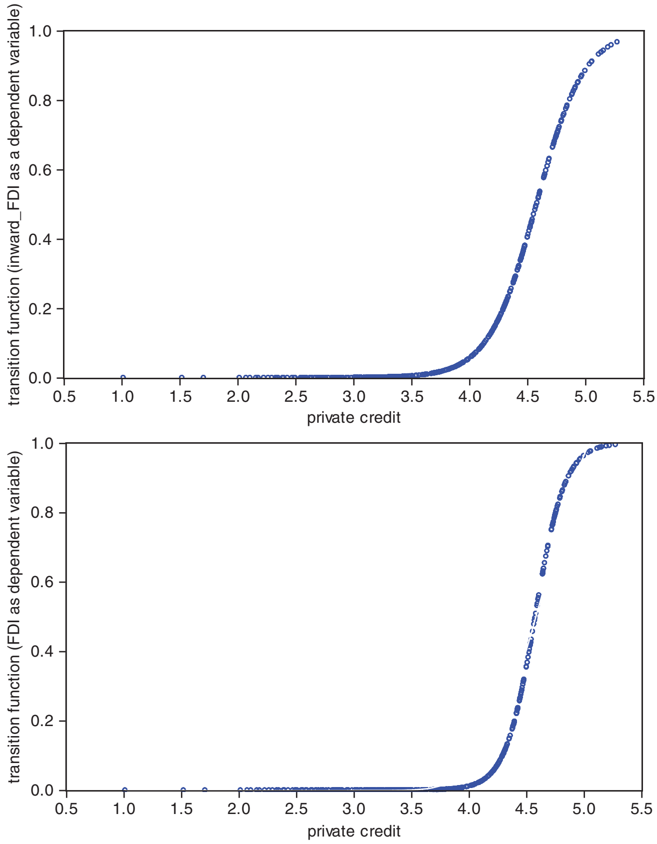
11. It is also noted that the estimation result for the direct impact of financial development in Model 3 indicates that financial development widens income inequality at the lower regime of financial development but tends to improve the income distribution at the higher regime. However, both effects are not statistically significant. The findings here imply that the inequality impact of financial development is very likely through its interaction with FDI.

### *B. Robustness Checks*

As a robustness check, we experiment with an alternative financial development measure: private credit. The results are reported in Table 4. We find results similar to those above. As Panel A indicates, the null hypothesis of linearity is strongly rejected, and the optimal number of transition functions, $r$, is 1. According to Panel B of Table 4, the PSTR model with $m=1$ is chosen when $F_1$ is larger than other statistics. Also, as illustrated in Panel C, the slope parameter is small, indicating the transition function is a smooth and continuous function of financial development. This result is shown in Figure 2 for the case of Model 3 in Table 4. Both models have a property of the smallest slope parameter. Furthermore, the coefficient estimates $\widehat{\beta}_0$ and $\widehat{\beta}_1$ remain positive and statistically significant, regardless of different model specifications and FDI indicators.

**FIGURE 2**

Estimated Transition Function of the PSTR Model against Financial Development for Model 3 of Table 6.




*Note:* $y$ axis is the transition function $g(q_{it}; \gamma, c)$ and the $x$ axis is the transition variable: private credit.

Therefore, as FDI increases, income inequality slowly increases in the early stage of financial development and then rapidly increases in the later stage after the levels of financial development exceed the threshold value of 3.3917–4.5790.

As another robustness check, Table 5 considers an alternative inequality measure (*gini_gross*) for the full set of control variables. Similar results are found. As Panel A indicates, the null hypothesis of linearity is strongly rejected and the optimal number of transition functions is 1. In Panel B, the PSTR model with $m = 1$ is chosen as $F_1$ is larger than other statistics. Also, as illustrated in Panel C, the slope parameter is small, indicating the transition function is a smooth and continuous function of financial development. The coefficient estimates $\widehat{\beta}_0$ and $\widehat{\beta}_1$ remain positive and statistically significant, regardless of different inequality measures and FDI indicators. FDI increases income inequality, and this effect increases with financial development.

As argued by Figini and Görg (2011), outward FDI should bring, in the host country, the opposite effect of inward FDI, since it is related with technology transfers abroad. Bitzer and Görg (2009) indeed find inward FDI is positively associated with domestic productivity at the industry level, while this relationship is negative for outward FDI. It is thus interesting to examine whether outward FDI shares the same pattern as inward FDI. Table 6 reports the estimation results. As expected, the optimal PSTR model for outward FDI is when $r = 1$ and $m = 1$ (Panels A and B), where the slope parameter is small, indicating the transition function is a smooth

**TABLE 5**
Parameter Estimates for the Final PSTR Models

| Dependent variable: *gini_gross* Threshold variable | *Findev* | | *Private Credit* | |
|---|---|---|---|---|
| | **Inward FDI** | **Total FDI** | **Inward FDI** | **Total FDI** |
| Panel A: Tests for linearity and remaining nonlinearity | | | | |
| $m=1$ | | | | |
| $H_0: r=0$ vs. $H_1: r=1$ | 1.720* | 2.330** | 2.236** | 2.604*** |
| | [0.092] | [0.019] | [0.024] | [0.009] |
| $H_0: r=1$ vs. $H_1: r=2$ | 0.574 | 0.723 | 1.203 | 1.092 |
| | [0.799] | [0.672] | [0.296] | [0.368] |
| $m=2$ | | | | |
| $H_0: r=0$ vs. $H_1: r=1$ | 1.394 | 2.116*** | 2.479*** | 3.214*** |
| | [0.141] | [0.007] | [0.001] | [0.000] |
| $H_0: r=1$ vs. $H_1: r=2$ | 0.415 | 0.615 | 1.027 | 2.007 |
| | [0.979] | [0.872] | [0.427] | [0.012] |
| Panel B: Tests for choosing the PSTR model | | | | |
| $H_0^3: \theta_2 = 0$ | 0.99 | 0.92 | 1.76* | 1.80* |
| $F_3$ | [0.4460] | [0.4970] | [0.0828] | [0.0767] |
| $H_0^2: \theta_1 = 0 \mid \theta_2 = 0$ | 0.53 | 0.33 | 1.65 | 1.51 |
| $F_2$ | [0.8341] | [0.9548] | [0.1092] | [0.1537] |
| $H_0^1: \theta_0 = 0 \mid \theta_1 = \theta_2 = 0$ | 327.98*** | 330.70*** | 329.17*** | 330.03*** |
| $F_1$ | [0.0000] | [0.0000] | [0.0000] | [0.0000] |
| Panel C: Parameter estimates | | | | |
| Initial FDI: $\beta_0$ | 0.0064* | 0.0051* | 0.0098*** | 0.0050* |
| | (0.0034) | (0.0030) | (0.0034) | (0.0029) |
| $\beta_1$ | 0.0077* | 0.0110* | 0.0083* | 0.0066* |
| | (0.0047) | (0.0061) | (0.0048) | (0.0040) |
| Location parameters | 2.1839 | 2.3685 | 3.9924 | 3.4778 |
| Slopes parameters | 12.2369 | 11.5471 | 8.5508 | 19.0239 |
| SSR | 0.3739 | 0.3363 | 0.3664 | 0.3079 |
| AIC | −6.8921 | −6.9979 | −6.9123 | −7.0861 |
| BIC | −6.7190 | −6.8248 | −6.7391 | −6.9130 |
| Observations (N*T) | 420 | 420 | 420 | 420 |

*Note*: The standard errors ($p$ values) in parentheses (brackets) are corrected for heteroskedasticity.
***Significant at 1%; **significant at 5%; *significant at 10%.

and continuous function of financial development (Panel C). Interestingly, now the estimated coefficient $\widehat{\beta}_0$ turns negative while $\widehat{\beta}_1$ remain positive, and both are statistically significant, regardless of different model specifications and financial development indicators. Further, the significantly negative coefficient estimate $\left(\widehat{\beta}_0\right)$ and the significantly positive sum of $\widehat{\beta}_0$ and $\widehat{\beta}_1$ suggest that as outward FDI increases, income inequality decreases in the early stage of financial development and then increases in the later stage after the levels of financial development exceed the threshold value of 2.020−3.930. Outward FDI brings the opposite effect of inward FDI, especially at the early stages of financial development. The evidence hence indicates that while total FDI raises income inequality, outward FDI is the potential component of total FDI which improves income distribution at the early stages of financial development.

Given the parameter estimates of the final PSTR models, it is now possible to compute, for each country in the sample and for each period, the time-varying coefficient for the effect of FDI on income inequality, that is, $\frac{\partial \mathrm{ineq}_{it}}{\partial \mathrm{fdi}_{it-1}}, \forall i = 1, 2, 3, \ldots, N$, and $\forall t = 1, 2, 3, \ldots, T$. The individual averages of these smoothed coefficients for FDI, as well as their standard deviations, are reported in Table 7. These estimates are based on the historical values of the transition variable $q_{it-1}$ observed for the 42 countries, that is, the mean by country of the individual estimates, $B_i = \frac{1}{T}\sum_{t=1}^{T} \frac{\partial \mathrm{ineq}_{it}}{\partial \mathrm{fdi}_{it-1}}$. To compare our model with linear specifications, we also include the results obtained in time series and in a linear homogenous panel with fixed individual effects. Consider the PSTR estimates derived from Model 3 of Table 3 for the case of inward FDI which is shown in the first column. The average estimated impact of

**TABLE 6**
PSTR Estimates for Outward FDI

Dependent variable: *gini_net*
Threshold variable: *findev* (*models* 1 & 2) *and private credit* (*models* 3 & 4)
Panel A: $LM_F$ tests for remaining nonlinearity

| | $m=1$ | | | | $m=2$ | | | |
|---|---|---|---|---|---|---|---|---|
| | **Model 1** | **Model 2** | **Model 3** | **Model 4** | **Model 1** | **Model 2** | **Model 3** | **Model 4** |
| $H_0: r=0$ vs. $H_1: r=1$ | 2.154* | 3.029*** | 2.476* | 2.482** | 2.179** | 2.571*** | 2.111* | 2.504*** |
| | (0.094) | (0.003) | [0.062] | [0.024] | (0.046) | (0.001) | [0.053] | [0.004] |
| $H_0: r=1$ vs. $H_1: r=2$ | 1.456 | 0.644 | 0.800 | 0.943 | 1.032 | 0.378 | 0.638 | 1.579* |
| | (0.228) | (0.740) | [0.495] | [0.465] | (0.405) | (0.986) | [0.700] | [0.100] |

Panel B: Tests for choosing the PSTR model for the period 1976–2005

| $r^*=1$ | **Model 1** | **Model 2** | **Model 3** | **Model 4** |
|---|---|---|---|---|
| $H_0^3: \theta_3 = 0$ | 4.18*** | 1.87* | 1.36 | 2.34** |
| $F_3$ | (0.0067) | (0.0660) | [0.2555] | [0.0330] |
| $H_0^1: \theta_2 = 0 \mid \theta_3 = 0$ | 6.33*** | 2.98*** | 2.39* | 2.41** |
| $F_2$ | (0.0004) | (0.0035) | [0.0698] | [0.0283] |
| $H_0^1: \theta_1 = 0 \mid \theta_2 = \theta_3 = 0$ | 851.84*** | 320.46*** | 855.46*** | 430.15*** |
| $F_1$ | (0.0000) | (0.0000) | [0.0000] | [0.0000] |

Panel C: Parameter estimates for the final PSTR models

| | **Model 1** | **Model 2** | **Model 3** | **Model 4** |
|---|---|---|---|---|
| Initial FDI: $\beta_0$ | −0.0096* | −0.0103*** | −0.0152*** | −0.0126* |
| | (0.0051) | (0.0041) | (0.0056) | (0.0076) |
| $\beta_1$ | 0.0211*** | 0.0164*** | 0.0261*** | 0.0308** |
| | (0.0056) | (0.0045) | (0.0067) | (0.0143) |
| Location parameters c | 2.1347 | 2.0206 | 3.3013 | 3.9302 |
| Slopes parameters | 6.2216 | 9.5626 | 2.6837 | 1.0037 |
| Sum of squared residuals | 0.079 | 0.071 | 0.0809 | 0.077 |
| AIC | −7.9185 | −7.8937 | −7.8905 | −7.8608 |
| BIC | −7.8025 | −7.6327 | −7.7745 | −7.6578 |
| Observations (N*T) | 240 | 240 | 240 | 240 |

*Notes*: Control variables in Models 1 and 3 include initial inequality and initial real per-capita GDP growth for each period. Models 2 and 4 include all control variables. The standard errors ($p$ values) in parentheses (brackets) are corrected for heteroskedasticity.

***Significant at 1%; **significant at 5%; *significant at 10%.

FDI on income inequality differs significantly between countries. The estimated coefficient of FDI is 0.7066 for Hungary, but 1.9830 for Japan. Likewise, in considering the PSTR estimates obtained from Model 3 of Table 3 for the case of total FDI which is shown in the second column, the estimated coefficient of FDI is 0.3781 for Hungary, but 1.5981 for Japan. These results suggest that our sample is highly heterogeneous and that this variation is due to differences in levels of financial development across countries. This finding is clearly illustrated by Figure 3, which displays the country average impact of FDI against the country average level of financial development. Countries with higher financial development tend to have a greater impact of FDI on income inequality.When comparing the results of a homogenous linear panel model with those of the PSTR model, we find large differences between them. For example, the average estimated coefficient of inward (total) FDI is 0.9619 (0.6678) in the PSTR model, whereas it is only 0.0109 (0.0065) in the panel model, as shown in the last row of Table 7. Utilizing time series may reduce the information set and produce unrealistic estimates (i.e., negative or too small) of the effect of FDI on income inequality. This result clearly illustrates the advantage of the PSTR approach.

The PSTR specification also allows us to study the time dynamics of the estimated effect of FDI on inequality over the period 1976–2005. Accordingly, we only consider Model 3 of Table 3, that is, the model with a full set of controls and total FDI as a dependent variable. The individual estimates for FDI parameters are displayed in Figure 4. These estimates are derived from Equation (4), but the difference between this case and the previous one (Figure 3) is that here we consider the historical value of the threshold variable $q_{it-1}$ observed for each country between 1976 and 2005. For most of

**TABLE 7**
Individual Estimates of the Effect of FDI on Inequality

| | | Inward FDI | | Total FDI | |
|---|---|---|---|---|---|
| **Country** | ***findev*** | **OLS/within B (*SD*)** | **PSTR B (*SD*)** | **OLS/within B (*SD*)** | **PSTR B (*SD*)** |
| Australia | 2.4281 | 0.0561 (0.0445) | 0.9107 (0.2094) | 0.0765 (0.0293) | 0.6552 (0.2083) |
| Belgium | 2.5324 | 0.0205 (0.0262) | 1.0763 (0.3944) | 0.0250 (0.0318) | 0.7816 (0.3316) |
| Canada | 2.6908 | 0.1282 (0.0111) | 1.2052 (0.3320) | 0.2679 (0.0007) | 0.9271 (0.2746) |
| Colombia | 2.1660 | −0.0293 (0.0245) | 0.7301 (0.0210) | −0.0281 (0.0255) | 0.4507 (0.0447) |
| Costa Rica | 1.7192 | −0.0651 (0.0748) | 0.7049 (0.0098) | −0.0616 (0.0759) | 0.3809 (0.0301) |
| Germany | 2.7553 | −0.0296 (0.0455). | 1.3315 (0.2830) | −0.0293 (0.1114) | 1.0261 (0.2083) |
| Denmark | 2.3196 | −0.0083 (0.0166) | 0.8917 (0.2040) | −0.0100 (0.0156) | 0.6286 (0.2169) |
| Dominican | 2.0224 | 0.0745 (0.0424) | 0.7114 (0.0099) | 0.0040 (0.0516) | 0.4066 (0.0272) |
| Ecuador | 1.7034 | 0.0040 (0.0010) | 0.7065 (0.0156) | 0.0044 (0.0011) | 0.3838 (0.0403) |
| Egypt | 1.8601 | 0.0012 (0.0233) | 0.7106 (0.0164) | 0.0005 (0.0232) | 0.3982 (0.0434) |
| Spain | 2.6430 | 0.0250 (0.1223) | 1.1458 (0.2609) | 0.0540 (0.0692) | 0.8775 (0.2002) |
| Finland | 2.5100 | 0.0127 (0.0053) | 0.9329 (0.1337) | 0.0148 (0.0056) | 0.7027 (0.1265) |
| France | 2.6828 | −0.0802 (0.1353) | 1.1985 (0.0969) | 0.0704 (0.1454) | 0.9280 (0.0735) |
| United Kingdom | 2.6461 | −0.0422 (0.0836) | 1.2593 (0.4792) | −0.0405 (0.0410) | 0.9295 (0.4049) |
| Greece | 2.1636 | −0.0215 (0.0057) | 0.7513 (0.0879) | −0.0415 (0.0218) | 0.4692 (0.1165) |
| Hungary | 1.2463 | 0.1385 (0.1518). | 0.7066 (0.0199) | −0.0759 (0.0803) | 0.3781 (0.0488) |
| Indonesia | 1.9839 | −0.0350 (0.1069) | 0.7431 (0.0711) | −0.0313 (0.1195) | 0.4502 (0.1149) |
| Ireland | 2.5173 | 0.0529 (0.0095) | 1.0241 (0.3339) | 0.0577 (0.0039) | 0.7512 (0.2813) |
| Israel | 2.4797 | 0.0067 (0.0011) | 0.9894 (0.3174) | 0.0111 (0.0009) | 0.7187 (0.2764) |
| Italy | 2.3968 | 0.0336 (0.0867) | 0.8187 (0.0759) | 0.0660 (0.1053) | 0.5843 (0.0876) |
| Jamaica | 1.8293 | −0.0310 (0.1618) | 0.7080 (0.0108) | −0.0829 (0.1883) | 0.3919 (0.0342) |
| Jordan | 2.5415 | 0.0474 (0.0226) | 1.0186 (0.2645) | 0.0465 (0.0214) | 0.7618 (0.2259) |
| Japan | 3.5960 | 0.0075 (0.0656) | 1.9830 (0.0261) | 0.1362 (0.0257) | 1.5981 (0.0490) |
| Korea | 2.4411 | −0.0494 (0.0034) | 0.9575 (0.3114) | −0.0567 (0.0199) | 0.6862 (0.2739) |
| Sri Lanka | 1.6740 | −0.0475 (0.0115) | 0.7075 (0.0129) | −0.0481 (0.0128) | 0.3872 (0.0399) |
| Mexico | 1.9498 | −0.0023 (0.0091). | 0.7152 (0.0219) | −0.0024 (0.0091). | 0.4101 (0.0514) |
| Malaysia | 2.8770 | 0.0172 (0.0519) | 1.5566 (0.4015) | 0.0167 (0.0353) | 1.1836 (0.2936) |
| Netherlands | 2.8293 | 0.0154 (0.1904) | 1.4746 (0.3371) | −0.0388 (0.2445) | 1.1286 (0.2487) |
| Norway | 2.5661 | 0.1234 (0.1634) | 0.9987 (0.1209) | −0.0711 (0.0552) | 0.7668 (0.1077) |
| Nepal | 1.8663 | 0.0001 (0.0588) | 0.7124 (0.0271) | 0.0234 (0.0301) | 0.4191 (0.1237) |
| New Zealand | 2.4437 | −0.0178 (0.0081) | 1.0977 (0.4019) | −0.0187 (0.0089) | 0.7793 (0.3670) |
| Pakistan | 1.8355 | 0.1774 (0.0170) | 0.7082 (0.0190) | −0.0504 (0.0413) | 0.3890 (0.0449) |
| Panama | 2.1015 | 0.0216 (0.0334) | 0.7562 (0.0927) | 0.0354 (0.0015) | 0.4699 (0.1314) |
| Philippines | 2.1138 | 0.0142 (0.0055) | 0.7302 (0.0339) | 0.0264 (0.0241) | 0.4436 (0.0683) |
| Portugal | 2.6876 | −0.0357 (0.0357) | 1.2490 (0.4080) | −0.0193 (0.0260) | 0.9458 (0.3197) |
| El Salvador | 1.8479 | −0.0228 (0.0074). | 0.7131 (0.0196) | 0.0291 (0.0467). | 0.4022 (0.0525) |
| Sweden | 2.4333 | 0.1011 (0.0562) | 0.8417 (0.0812) | −0.0232 (0.0075) | 0.6139 (0.0824) |
| Thailand | 2.6214 | 0.0271 (0.0063) | 1.2400 (0.5057) | −0.0564 (0.0542) | 0.9099 (0.4287) |
| Turkey | 1.8353 | 0.0088 (0.0194) | 0.7097 (0.0143) | 0.0070 (0.0097) | 0.3963 (0.0399) |
| United States | 2.6803 | 0.3043 (0.0996) | 1.1936 (0.1196) | −0.0251 (0.0538) | 0.9241 (0.0872) |
| Venezuela | 2.0429 | 0.0323 (0.0172) | 0.8418 (0.3299) | −0.1582 (0.2944) | 0.4852 (0.1368) |
| South Africa | 2.5371 | 0.2247 (0.0957) | 0.9394 (0.1447) | 0.0463 (0.1093) | 0.7285 (0.1050) |
| All Country | 2.3052 | 0.0109 (0.0513) | 0.9619 (0.1562) | 0.0065 (0.0661) | 0.6678 (0.1158) |

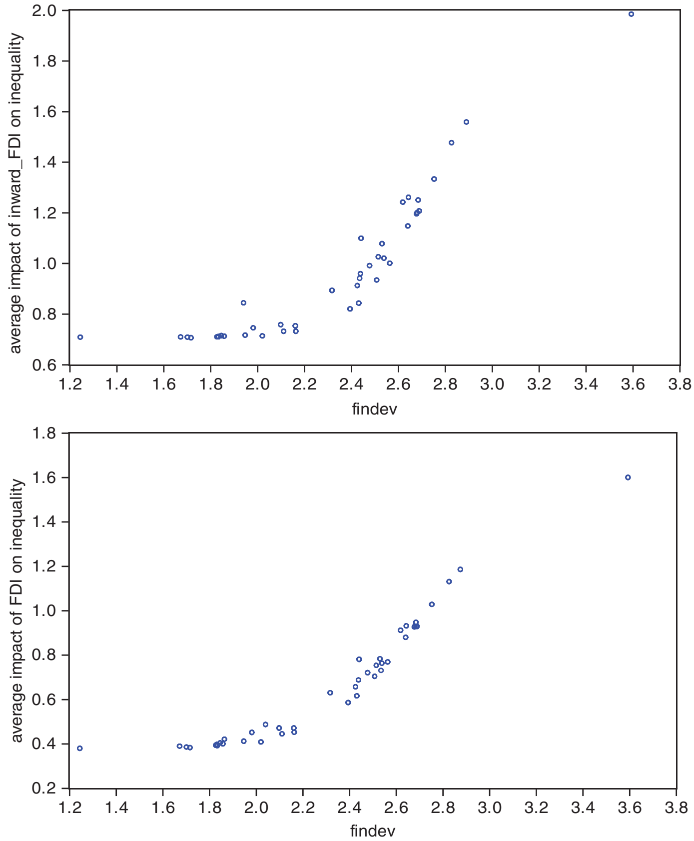
*Note*: The dependent variable is *gini_net*. For each country, the average estimated impact B and the standard deviation (in percentages) over the total period are reported. The first and third columns of the line for all countries correspond to the within estimates.

the countries in our sample, the estimates of FDI increase between 1976 and 2005. However, we can observe that the increase is generally moderate except for Belgium, New Zealand, Thailand, and the United Kingdom.

## V. CONCLUSIONS

This article empirically examines whether financial development influences the impact of FDI on income inequality. We employ the PSTR methodology that provides a more intuitive and flexible framework to deal with both cross-country heterogeneity and parameter stability, the two very typical issues to researchers in panel studies. Using a panel of advanced and developing countries over the period 1976–2005, we find that the impact of FDI—both inward and total—on income inequality is nonlinear and changes over time and across countries depending on the level of financial development. More specifically,

**FIGURE 3**
Average Estimated Impact of FDI on Income Inequality across Countries Based on Models 3 and 5 of Table 3.




*Note:* For each country, the observation represents the average estimated impact over the total period under consideration against the corresponding average level of financial development index.

FDI increases income inequality, and this effect strengthens with financial development. In other words, FDI raises income inequality slowly in a low-financial-development regime, but rapidly in a high-financial-development one. As expected, we find that outward FDI brings the opposite effect of inward FDI, particularly at the early stages of financial development. As outward FDI increases, income inequality decreases in the early stage of financial development and then increases in the later stage once financial development crosses some critical levels.

The evidence of inequality-enhancing effect of FDI is consistent with the arguments that FDI benefits the wealthy and increases demand for skilled and educated workers as the transferred technology are more capital- or skill-intensive (e.g., Feenstra and Hanson 1996). Furthermore, the finding that financial development reinforces the positive effect of FDI on income inequality seems in agreement with the argument put forward by Easterly (2001) and Milanovic (2003) that FDI reforms often lead to established, well-connected individuals capturing much of the gains from the new opportunities, especially when openness to FDI is coupled with market-oriented policy reforms such as the liberalization of the domestic labor market or the privatization of state-owned firms. Here, it is financial sector liberalization that increases inequality in income distribution when interacting with FDI openness, perhaps because the role of financial development as a precondition

**FIGURE 4**

Individual estimated impact of FDI on income inequality. The line represents the estimated impact of FDI on income inequality using the historical value of the threshold variable *findev*.

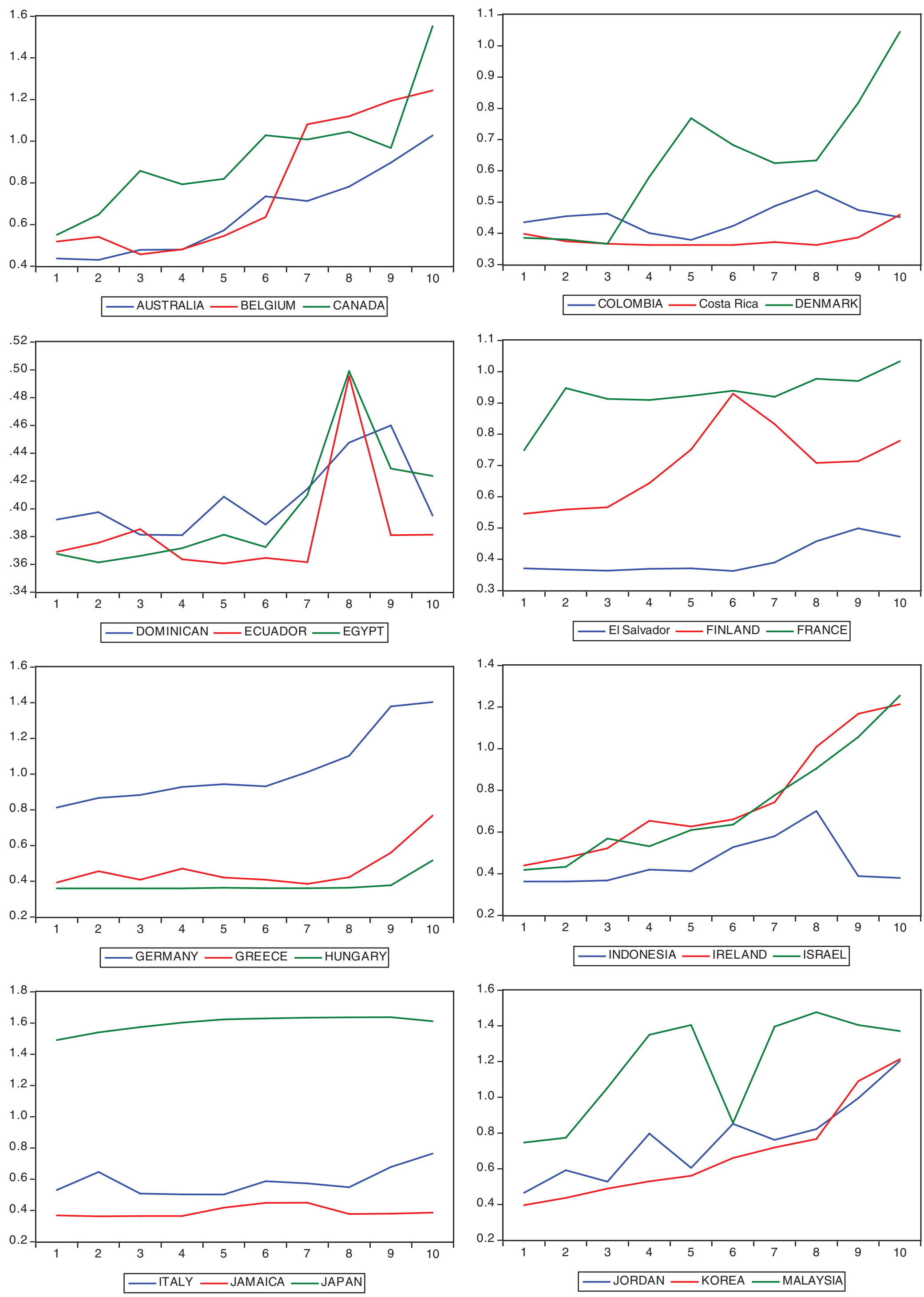
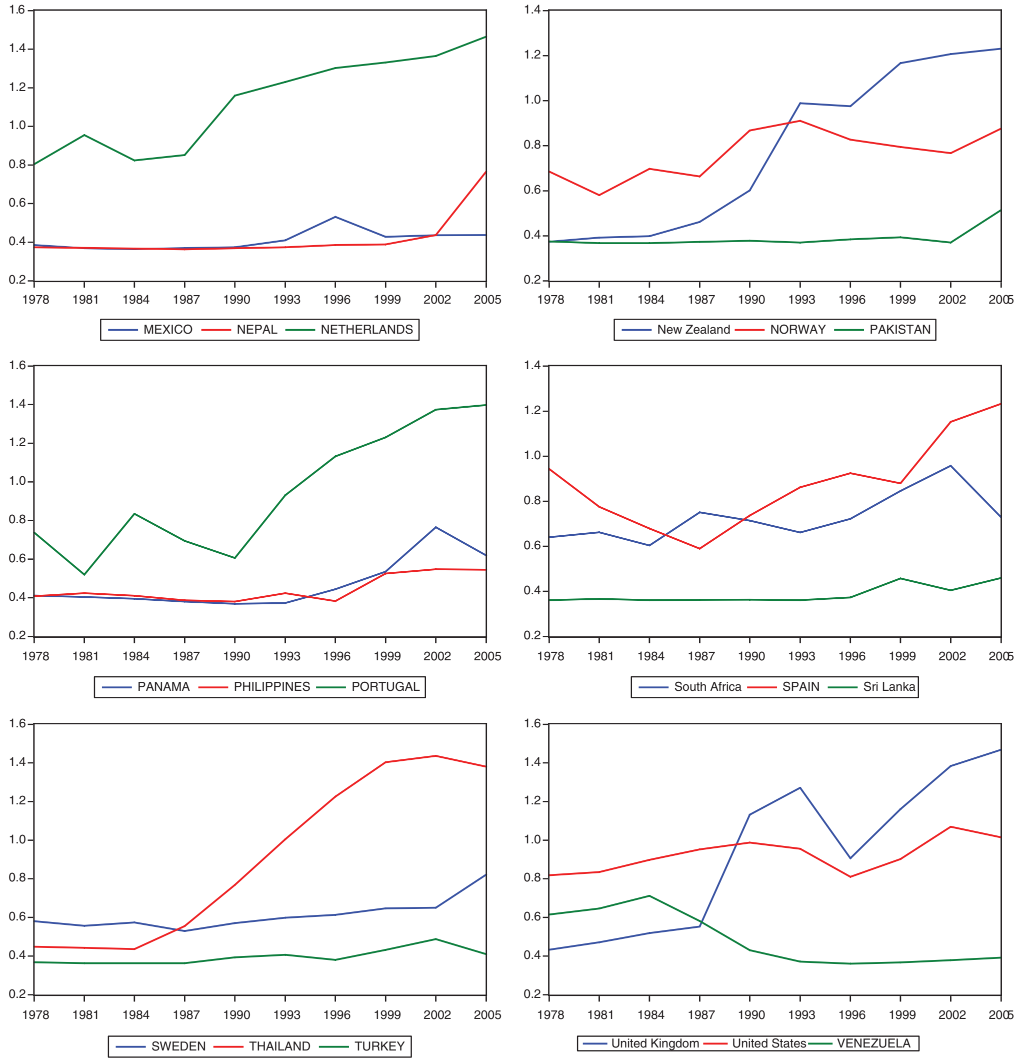


**FIGURE 4**
Continued.




to reap the growth benefits of FDI spillovers, according to the absorptive capacity hypothesis, but at the expense of the poor following the new political economy considerations. Ideally, financial development makes it possible for much of the population to realize profit opportunity from FDI by increasing the availability for previously credit constrained individuals to access capital. In actuality, financial development enhances the financial services of those already accessing the financial system. They are often high-income individuals or well-established firms. Thus, the direct effect from improving the quality of financial services falls disproportionately on the rich, widening inequality and perpetuating cross-dynasty differences in economic opportunities from FDI.

Our finding thus has an important policy implication. In order to expand disadvantage groups' economic opportunities associated with FDI liberalization, how to avoid financial sector reform policy to be captured and controlled by the rich and political elites should be high on the agenda for governments that intend to improve economic growth with a more egalitarian income distribution through FDI.

APPENDIX

**TABLE A1**
Descriptive Statistics

| | *Gini_net* | *Initial gini_net* | *Total fdi* | *Inward fdi* | *Private credit* | *Findev* | *Schooling* | *Rgdppc_gr* | *Inflation* | *Trade* | *Gov_spending* |
|---|---|---|---|---|---|---|---|---|---|---|---|
| Panel A: Summary statistics | | | | | | | | | | | |
| Mean | 3.5468 | 3.5443 | 2.6543 | 2.2651 | 3.7890 | 2.3051 | 2.2590 | 0.1046 | 4.0755 | 2.7196 | 1.9596 |
| *SD* | 0.2586 | 0.2596 | 1.1697 | 1.1755 | 0.7462 | 0.5014 | 3.3303 | 0.1353 | 0.5090 | 0.3847 | 0.4137 |
| Maximum | 4.2534 | 4.2163 | 5.3753 | 4.9510 | 5.2745 | 4.1089 | 19.9187 | 1.3977 | 5.3955 | 3.7069 | 2.5517 |
| Minimum | 3.0115 | 3.0073 | −0.9429 | −1.0894 | 1.3124 | 0.6671 | −15.1356 | −0.1012 | 2.2085 | 1.1690 | −0.2533 |
| Panel B: Correlation matrix | | | | | | | | | | | |
| *Gini_net* | 1.0000 | | | | | | | | | | |
| *Initial gini_net* | 0.9968 | 1.0000 | | | | | | | | | |
| *Total fdi* | 0.0052 | 0.0026 | 1.0000 | | | | | | | | |
| *Inward fdi* | 0.2030 | 0.2015 | 0.9190 | 1.0000 | | | | | | | |
| *Private credit* | −0.3708 | −0.3764 | 0.4525 | 0.2079 | 1.0000 | | | | | | |
| *Findev* | −0.2919 | −0.2937 | 0.4103 | 0.1520 | 0.8467 | 1.0000 | | | | | |
| *Rgdppc_gr* | −0.1380 | −0.1293 | −0.0923 | −0.1103 | 0.0516 | 0.1294 | 1.0000 | | | | |
| *Inflation* | 0.2566 | 0.2549 | −0.2413 | −0.1308 | −0.4120 | −0.3811 | −0.2140 | 1.0000 | | | |
| *Trade* | −0.0157 | −0.0157 | 0.4793 | 0.5716 | 0.1176 | 0.0185 | 0.0744 | −0.1007 | 1.0000 | | |
| *Gov_spending* | −0.4910 | −0.4929 | 0.4131 | 0.2358 | 0.5057 | 0.4128 | −0.0467 | −0.1864 | 0.2158 | 1.0000 | |
| *Schooling* | −0.4023 | −0.4102 | 0.5325 | 0.3444 | 0.5601 | 0.4386 | −0.0136 | −0.1568 | 0.2033 | 0.4453 | 1.0000 |

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