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FDI and Income Inequality - Evidence from Latin American Economies

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Abstract:

We analyze whether foreign direct investment (FDI) has contributed to the typically wide income gaps in Latin America. We perform panel cointegration techniques as well as regression analysis to assess the impact of inward FDI stocks on income inequality among households in Latin American host countries. The panel cointegration analysis typically reveals a significant and positive effect on income inequality. There is no evidence for reverse causality. Our findings are fairly robust to the choice of different estimation methods, sample selection, and the period of observation.

Keywords: FDI, income inequality, cointegration techniques, Latin America.

JEL classification: F21; D31

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1. Introduction

Latin America stands out as “the most economically unequal region in the world.”¹ Recent trends reveal, however, that income inequality has declined throughout the region – which is in striking contrast to widening income gaps in Asia, notably in China and India (López-Calva and Lustig 2010; Gasparini and Lustig 2011). At the same time, Latin America reported a stronger increase in foreign direct investment (FDI) than developing Asia since the 1990s. UNCTAD data reveal that inward FDI stocks in Latin America were less than one third of Asia’s inward FDI stocks in 1990. During the 2000-2011 period, Latin America hosted FDI in the order of half the Asian FDI stock. Measuring FDI as a percentage of GDP, Latin America became even more attractive than Asia.²

Conventional wisdom suggests that recent trends in inequality and FDI might support economic growth in Latin America. Several studies have found that higher inequality tends to retard growth in developing countries (Barro 2000), even though the empirical evidence is far from conclusive.³ FDI is widely believed to spur economic growth in the host countries as it brings superior technologies and know-how in addition to foreign capital (e.g., OECD 2002). Even globalization critics, including Stiglitz (2000), find the case for FDI compelling.⁴

Against this backdrop, it is not surprising that income redistribution (e.g., through poverty reduction programs) as well as FDI promotion figure high on the agenda of policymakers in Latin America. It has received only scant attention that this agenda may involve a dilemma. Specifically, the promotion of inward FDI may undermine efforts to narrow income gaps through

¹ <http://justf.org/blog/2010/06/08/income-inequality-latin-america-today> (accessed: August 2012). See also World Bank (2004).

² For details see: <http://unctadstat.unctad.org/> (accessed: August 2012).

³ Banerjee and Duflo (2003) argue that “efforts to interpret this evidence causally run into difficult identification problems.” Klasen and Lamanna (2009) focus on gender inequality, finding that gender gaps in education and employment considerably reduce economic growth. Grimm (2011) investigates the effects of inequality in health on economic growth, finding a substantial and adverse effect in low and middle income countries.

⁴ However, Alfaro et al. (2010) conclude from the recent empirical literature that the macroeconomic evidence for positive growth effects of FDI in developing countries continues to be weak.

redistribution if FDI leads to greater inequality in the host country. As we discuss in Section 2, the relationship between FDI and income inequality is theoretically ambiguous. Moreover, previous empirical evidence for developed host countries, notably the United States, does not necessarily hold for less advanced Latin American host countries.

Therefore, we perform panel cointegration analyses to assess the distributional effects of inward FDI in Latin American countries since the early 1980s. Following the discussion of the theoretical background in Section 2, we present the empirical model and the data used in Section 3. We report the estimation results in Section 4. We find that higher inward FDI stocks typically widen the income gaps in Latin American host countries. Section 5 summarizes and concludes.

2. Theoretical background and previous findings

The theoretical literature on inward FDI departs from the observation that multinational enterprises (MNEs) possess firm-specific assets such as technological knowledge and management skills, granting them a productivity advantage over domestic firms in the host country. The heterogeneous firm model of Helpman et al. (2004) predicts that only the most productive firms engage in FDI to serve foreign markets. Ownership advantages are required to overcome the ‘liability of foreignness’, i.e., the lacking familiarity with conducting operations in the home market of local firms (Markusen 1995; Dunning and Lundan 2008).

It is consistent with the productivity advantages of MNEs that they are generally found to pay higher wages than local firms (Aitken et al. 1996; Lipsey 2002). More specifically, MNEs may pay higher wages to discourage worker turnover.⁵ Importantly, a review of the empirical literature reveals that “almost all evidence shows that FDI and foreign ownership are associated

⁵ MNEs have an incentive to reduce worker turnover as they incur higher search costs than domestic firms which are familiar with local labor markets. Furthermore, MNEs tend to invest more in training. Higher wages may also help contain the leakage of firm-specific assets to domestic firms.

with higher wages for *all* types of workers” (Overseas Development Institute 2002: 2; emphasis added).

This evidence suggests that the fierce competition for FDI among potential host countries in Latin America and elsewhere does not necessarily undermine efforts at reducing income inequality. FDI would even support such efforts in a Heckscher-Ohlin framework. In such a framework, FDI inflows resemble trade liberalization in that the relatively abundant factor of production would benefit. Latin America is often assumed to be abundant in less skilled labor (Robertson 2000). By contrast, more advanced countries with an abundant supply of skilled labor are the principal sources of FDI in Latin America. Consequently, FDI from advanced countries in Latin America would increase income inequality in the source countries and reduce income inequality in the host countries.

Theoretical predictions become more complex when refining the ranking of skill intensities. Sorting MNE activities by skill intensity, Markusen and Venables (1997) consider headquarter (HQ) services to be more skill intensive than plant operations by MNEs. Domestic firms producing for the local market are least skill intensive and rank at the bottom of this classification. It has also to be taken into account that countries hosting plant operations by foreign MNEs may, at the same time, be home to HQ services of domestic MNEs. The establishment of foreign plant operations through FDI may then reduce the relative demand for skilled labor in the host country. This is most likely to happen where the HQ services of various domestic MNEs have traditionally shaped the demand for skilled labor. Inward FDI in the United States may be the most obvious case in point (Blonigen and Slaughter 2001).⁶ Low-income countries lacking HQ services of domestic MNEs tend to be at the other end of the spectrum of

⁶ Blonigen and Slaughter (2001) do not find any evidence that inward FDI contributed to skill upgrading in US manufacturing until the mid-1990s. Chintrakarn et al. (2012) perform panel co-integration analyses for US states, finding that FDI at the state level reduced income inequality during the 1977-2001 period.

host countries; for them inward FDI is most likely to increase the average skill intensity of production. Latin American host countries range in the middle ground. Several countries in the region increasingly emerged as home bases of domestic MNEs in the more recent past (Chudnovsky and López 2000; UNCTAD 2006; Santiso 2007). Theoretical predictions on the distributional effects of inward FDI become more ambiguous under such conditions.

FDI relations among similarly advanced source and host countries are predominantly of the horizontal type (Markusen 1995).⁷ By contrast, North-South models along the lines of Feenstra and Hanson (1997) focus on vertical FDI relations between more advanced source countries in the North and less advanced host countries in the South. Vertical FDI involves the fragmentation of production and provides a means to allocate specific steps of the production process to where the relevant comparative advantages can be utilized.⁸ Investors make use of varying factor endowments and differences in factor prices across countries (Markusen and Zhang 1999).

North-South models of vertical FDI figured most prominently in the context of the formation of the North American Free Trade Agreement (NAFTA). The availability of relatively cheap labor in Mexico and its proximity to US markets encouraged MNEs based in more advanced source countries, notably in the United States, to undertake vertical FDI by offshoring labor intensive parts of the production process to Mexico. According to Feenstra and Hanson (1997), this type of FDI may adversely affect the wage and employment prospects of less skilled workers not only in the advanced source countries, but also in the less advanced host country. This could happen if offshoring involves activities that are relatively skilled-labor intensive in the host country, even though they are relatively unskilled-labor intensive by the standards of the

⁷ Horizontal FDI is motivated by the attractiveness of host-country markets; MNEs duplicate the parent company's production at home in the host countries of FDI. For an early model of horizontal FDI, see Markusen (1984); more recent models include Markusen and Venables (1998; 2000).

⁸ For an early model of vertical FDI, see Helpman (1984).

source country. In contrast to the traditional Heckscher-Ohlin framework, inward FDI would then widen wage inequality in developing host countries.⁹

Several empirical studies support the hypothesis that FDI is associated with greater inequality by raising the skill premium in poorer host countries. For instance, inward FDI has benefited skilled workers more than unskilled workers in some Asian emerging economies, including Indonesia (Lipse and Sjöholm 2004), Korea (Mah 2002), and Thailand (te Velde and Morrissey 2004).¹⁰ As noted before, Mexico has received particular attention among Latin American host countries (e.g., Aitken et al. 1996; Feenstra and Hanson 1997). Hanson (2003) concludes from a survey of the earlier literature that FDI (and trade liberalization) has increased the relative demand for skilled labor in Mexico.

It remains open to question, however, whether the findings for Mexico are representative of Latin America. While Mexico has attracted vertical FDI in the context of NAFTA, horizontal FDI may play a more important role in other Latin American host countries. Das (2002) argues that the predictions of the model of Feenstra and Hanson (1997) critically depend on the assumption of free trade. Under free-trade conditions the developing host country would specialize in relatively unskilled-labor intensive production so that “capital movement to the South from the North takes place in the relatively skilled labor intensive stages of production at the margin, pushing the relative wage up” (Das 2002: 71). This scenario is most reasonable in the context of NAFTA. Other parts of Latin America appear to be “incompletely specialized”,

⁹ It should be noted, however, that Das (2002) comes to the opposite conclusion. Two factors contribute to the FDI-induced reduction in relative wages in Das’ theoretical model: First, foreign firms operating with superior technology in skilled-labor intensive sectors of developing economies gain market shares at the expense of less efficient domestic firms in these sectors. This shift in output to more efficient foreign firms involves some savings in terms of factor use, which mainly affects skilled labor in skilled-labor intensive sectors. The weaker relative demand for skilled labor reduces the relative wage. Second, the entry of more efficient foreign firms tends to increase the supply of skilled workers. This is because skilled local entrepreneurs are crowded out as owners and managers of domestic firms and join the labor force on which foreign firms can draw.

¹⁰ However, according to te Velde and Morrissey (2004), the effects of inward FDI on wage inequality are less clear or insignificant in Singapore, Hong Kong, the Philippines, and Korea.

however, due to remaining trade barriers. Hence, inward FDI would not necessarily take place in the relatively skilled-labor intensive stages of production. The relative wage effects of FDI are then harder to predict.

Finally, theoretical arguments suggest that the relationship between inward FDI and inequality is non-linear and varying over time once learning and skill upgrading in the “transition to a new technological paradigm” is taken into account (Aghion and Howitt 1998: 262). While domestic firms may benefit from FDI-induced spillovers, their absorption of new technologies may increase inequality in the short run and reduce inequality in the longer run. Aghion and Howitt (1998: chapter 8) model such a transition by explicitly referring to the Kuznets inverted-U hypothesis of rising and then falling inequality. Accordingly, the skill premium increases as long as learning efforts result in high demand for skills that are in short supply. FDI typically clusters in the host countries’ economic centers where skilled and mobile workers tend to be employed in dynamic, export-oriented and technologically advanced sectors. Income gaps are initially likely to widen as FDI induces growth mainly in these urban centers where the pool of sufficiently skilled labor is fixed in the short term. Subsequently, wage inequality declines to the extent that the supply of the required skills improves and firms have managed the transition to the new technological paradigm. Labor migration from the hinterland to the economic centers may help spreading the benefits of FDI-related spillovers and wage increases.¹¹

Drawing on the model of Aghion and Howitt (1998), Figini and Görg (1999: 596) regard MNEs “as ‘role models’ for indigenous firms.” Figini and Görg (1999) find evidence for transitional inequality in Ireland due to FDI-induced transfers of new technologies, know-how, and ideas. The Irish case reveals an inverted U-shaped pattern, with FDI first increasing and then

¹¹ However, a more equal distribution of FDI-related benefits depends on whether the growth effects are strong enough to induce worker mobility and on whether migrant workers from the hinterland are sufficiently qualified.

later reducing inequality.¹² It is open to debate, however, whether FDI-induced inequality is likely to be a transitional phenomenon in Latin America. According to Basu and Guariglia (2007), FDI-induced inequality may rather persist unless poor population segments are able to accumulate sufficient human capital required to handle modern technologies. Various studies reveal that human capital formation in Latin American countries lags considerably behind countries with similar average per-capita incomes in other regions (e.g., Arellano 2002; Puryear and Goodspeed 2008). Sachs and Vial (2002: 13) conclude from their assessment of Latin America's international competitiveness: "Low investment in human capital in the past has been compounded by today's low levels and poor yields of investment in education, affecting the ability of future generations of workers to innovate and integrate successfully into a knowledge-based economy."

Theoretical ambiguity calls for empirical research on the distributional effects of FDI. However, apart from the country-specific studies mentioned before, empirical studies focusing on low and middle income host countries are still few. Some indications exist that the distributional consequences of FDI in developing host countries differ from those in more advanced host countries (Gopinath and Chen 2003; Figini and Görg 2011). Yet, the cross-country evidence for developing countries is inconclusive. Tsai (1995: 480) reckons that statistically significant correlations between FDI and income inequality reflect structural differences in inequality between geographical country groups, rather than implying a "deleterious influence of FDI." By contrast, the cross-country study of Choi (2006) finds more pronounced income inequality where the ratio of FDI stocks to GDP is higher.¹³ The estimations of Basu and Guariglia (2007) for a

¹² Figini and Görg (2011) report two distinct patterns with regard to FDI-induced transitional inequality. Wage inequality initially widens with FDI in developing countries, while this effect diminishes with further increases in FDI. By contrast, non-linear effects do not play a significant role in advanced host countries of FDI.

¹³ Choi (2006) finds that income inequality is generally higher in Latin America than elsewhere, by including a dummy variable for Latin American host countries. However, Choi does not address the FDI-inequality link specifically for Latin America.

large sample of developing countries point to a trade-off between FDI-related growth promotion and rising inequality (in terms of schooling).

Previous empirical studies are often restricted to wage inequality in the manufacturing sector. This is an important limitation as FDI in the services sector has become increasingly important and may have different distributional effects. Furthermore, studies on relative wages and labor shares provide an incomplete picture on inequality, ignoring “self-employment income, property income, profits, and executive compensation” (Lindert and Williamson 2001: 34). We overcome these limitations by using data on broader inequality concepts available from the University of Texas Inequality Project as well as the Standardized World Income Inequality Database (SWIID). The subsequent cointegration analysis also addresses causality concerns that tend to impair earlier regression analyses.

3. Model and data

We analyze the relationship between income inequality and FDI in Latin America using panel cointegration techniques, in addition to more conventional regression techniques. Cointegration estimators are robust under cointegration to a variety of estimation problems that often plague empirical work, including omitted variables, endogeneity and measurement error. This section introduces the basic model, describes the data, and discusses some econometric issues.

Following Chintrakarn et al. (2012), we assume that the following bivariate equation is a correct specification of the long-run relationship between FDI and inequality:¹⁴

$$EHII_{it} = a_{1i} + a_2 \left(\frac{FDI}{GDP} \right)_{it} + e_{it}, \quad (1)$$

¹⁴ When not further specified, the term inequality refers to income inequality among households.

where EHI_{it} stands for the estimated household inequality in Gini format over time periods $t = 1, 2, \dots, T$ and countries $i = 1, 2, \dots, N$, and $(FDI/GDP)_{it}$ is the inward FDI stock relative to GDP. Following common practice (see, e.g., Figini and Görg 2011; Chintrakarn et al. 2012), we use FDI stocks rather than FDI flows because stocks capture long-run effects more effectively due to the accumulation of flows. By expressing the FDI stock as a percentage of GDP, we control for the size of the host country (as is also common practice). The coefficient a_2 measures the long-run effect of inward FDI on inequality, and the a_{1i} represent country-specific intercepts, capturing any country-specific omitted factors that are relatively stable over time.

Since the early 1980s, both inequality and FDI have increased sharply in most countries (see, e.g., Galbraith 2007). Hence, it is reasonable to assume that EHI_{it} and $(FDI/GDP)_{it}$ are nonstationary integrated processes. If this assumption is correct, the linear combination of these two variables must be stationary, or, in the terminology of Engle and Granger (1987), EHI_{it} and $(FDI/GDP)_{it}$ must be cointegrated. If the two variables are not cointegrated, there is no long-run relationship between inequality and FDI; Equation (1) would in this case be a spurious regression in the sense of Granger and Newbold (1974). As shown by Entorf (1997) and Kao (1999), the tendency for spuriously indicating a relationship may even be stronger in panel data regressions than in pure time series regressions. The requirement for the above regression not to be spurious is thus that the two (integrated) variables cointegrate.

If two or more variables are cointegrated, then the parameter estimates are superconsistent, meaning that they are not only consistent but converge to the true parameter values at a faster rate than is normally the case, namely rate T rather than \sqrt{T} (Stock 1987). Accordingly, we obtain more accurate estimates under cointegration than would be possible with conventional methods. As shown by Stock (1987), the estimated cointegration coefficients are

superconsistent even in the presence of temporal and/or contemporaneous correlation between the (stationary) error term and the regressor(s). Consequently, estimates of cointegrating relationships are not biased by omitted stationary variables.

The fact that a regression consisting of cointegrated variables has a stationary error term also implies that no relevant nonstationary variables are omitted. Any omitted nonstationary variable that is part of the cointegrating relationship would become part of the error term, thereby producing nonstationary residuals and thus leading to a failure to detect cointegration.

If there is cointegration between a set of variables, then this stationary relationship also exists in extended variable space. In other words, cointegration relationships are invariant to model extensions (Lütkepohl 2007). An important implication of finding cointegration is thus that no additional variables are required to produce unbiased parameter estimates.

Another econometric issue relates to the potential cross-country heterogeneity in the relationship between FDI and inequality. Latin American economies differ in terms of economic development, attractiveness to FDI and openness to trade to name just a few dimensions. Thus, we face a dilemma regarding the optimal estimation strategy. On the one hand, efficiency gains from the pooling of observations over the cross-sectional units can be achieved when the individual slope coefficients are the same. On the other hand, pooled within-dimension estimators produce inconsistent and potentially misleading point estimates of the sample mean of the heterogeneous cointegrating vectors when the true slope coefficients are heterogeneous (see, e.g., Pesaran and Smith 1995). Although a comparative study by Baltagi and Griffin (1997) concludes that the efficiency gains from pooling more than offset the biases due to individual country

heterogeneity, we try to solve this dilemma by using both homogeneous (within-dimension-based) and heterogeneous (between-dimension-based) estimators.¹⁵

We now describe the data used in our analysis. The FDI-to-GDP ratios are from the United Nations Conference on Trade and Development (UNCTAD) database (available at: <http://unctadstat.unctad.org>). FDI stocks comprise the value of the share of a company's capital and reserves that are attributable to the foreign parent company. This also includes intra-company loans.

Like earlier studies (e.g., Herzer and Vollmer 2012), we use the Estimated Household Income Inequality (EHII) dataset provided by the University of Texas Inequality Project (<http://utip.gov.utexas.edu/data.html>) in our baseline estimations. This dataset has the major advantage of being comprehensive and consistent. Comprehensiveness was achieved by combining information from the well-known Deininger and Squire (1996) inequality dataset with data on manufacturing pay dispersion and the rate of blue-collar employment to total population from the United Nations Industrial Development Organization (UNIDO). The detailed calculation methods of the EHII dataset are laid out in Galbraith and Kum (2005).

A potential disadvantage of panel cointegration methods is that they typically require balanced panel data over a sufficiently long time period. However, the data from this particularly reliable source are available only for a relatively small balanced panel of five Latin American countries covering the period from 1980 to 2000 (21 yearly observations per country).¹⁶ The time

¹⁵ We also ran country-specific regressions to examine the impact of FDI on inequality for each country individually. These estimations are not shown for the sake of brevity. They are available from the authors upon request.

¹⁶ Bolivia, Chile, Colombia, Mexico and Uruguay. Appendix table A1 provides summary statistics for the sample underlying our baseline estimations. Bolivia is the poorest country with the highest inequality in our sample. Mexico ranks at the bottom with the lowest income inequality among households. However, inequality in our sample is in general fairly high. Chile and Bolivia are the top FDI recipients. As noted above, the time series properties of our data on the EHII Gini coefficients and the FDI-to-GDP ratios appear to be consistent with the possibility that the series are nonstationary. This is confirmed by the Augmented Dickey-Fuller (ADF) tests reported in Appendix table A.2.

series are sufficiently long to conduct a cointegration analysis.¹⁷ In addition, we use several test and estimation methods to ensure the robustness of our results. Specifically, we use panel cointegration methods which have higher power (due to the exploitation of both the time-series and cross-sectional dimensions of the data) and therefore can be implemented with shorter data spans than their time-series counterparts. Nevertheless, we draw on alternative data sources in Section 4.2 in order to be able to cover the more recent past. In this way, we are also able to include a larger number of Latin American host countries.¹⁸

4. Empirical analysis

We first use panel cointegration techniques to examine the relationship between FDI and inequality for the small sample of five Latin American countries for which we have particularly reliable data from the preferred EHII dataset. Subsequently, we perform panel cointegration estimates for a larger Latin American sample and an extended period of observation by drawing on alternative data sources (Section 4.2). In Section 4.3, we turn to alternative techniques as unbalanced panel estimations allow us to further expand the country sample and cover the most recent past. Finally, we focus on more recent years in Section 4.4 to assess at least tentatively whether the FDI-inequality link has changed over time.

4.1. Baseline panel cointegration results

4.1.1. Panel cointegration tests

Before we start with estimating the long-run relationship given by Equation (1), we run the necessary pre-tests for cointegration. As discussed above, an advantage of panel cointegration

¹⁷ Several cointegration analyses for individual countries are based on shorter periods (e.g., Crombrughe et al. 1997; Irvin and Izurieta 2000).

¹⁸ For details see Section 4.2.

procedures is that their implementation is possible for shorter time periods compared to pure time series applications.

We use several panel cointegration test procedures to determine whether there is a long-run relationship between FDI and inequality in Latin America. The first is the two-step residual-based procedure suggested by Pedroni (1999, 2004), which can be intuitively described as follows. In the first step, the hypothesized cointegrating regression (Equation (1)) is estimated separately for each country, thus allowing for heterogeneous cointegrating vectors. In the second step, the residuals, $\hat{\epsilon}_{it}$, from these regressions are tested for stationarity. To test the null hypothesis of non-stationarity (or no cointegration) Pedroni proposes seven statistics. Here, we employ the four statistics with the highest power for small T -panels like ours: the panel and group ADF and PP test statistics (see, e.g., Pedroni 2004). The panel statistics pool the autoregressive coefficients across different countries during the unit root test on the residuals of the static cointegrating regressions, whereas the group statistics are based on averaging the individually estimated autoregressive coefficients for each country. The panel ADF statistic is analogous to the Levin et al. (2002) panel unit root test. The group ADF statistic is analogous to the Im et al. (2003) panel unit root test. The PP statistics are panel versions of the Phillips-Perron (PP) t -statistics.

In addition, we use the panel cointegration tests developed by Kao (1999). Kao follows basically the same approach as Pedroni (1999, 2004), but constrains the cointegrating coefficients to be homogeneous across countries. To test for the stationarity of the residuals, Kao presents four (within-dimension-based) DF test statistics and one within-dimension-based ADF statistic: The first two DF statistics, DF_{ρ} and DF_t , as well as the ADF statistic, assume strict exogeneity of the regressors, while the other two DF-type tests, DF_{ρ}^* and DF_t^* , do not require this

assumption. DF_{ρ} and DF_{ρ}^* are calculated based on the estimated first-order autoregressive coefficient in the panel DF regression; the associated t -statistic is used in calculating DF_t and DF_t^* .

The results of the cointegration tests are presented in Table 1. All test statistics reject the null hypothesis of no cointegration at the one percent significance level, suggesting that there is a long-run relationship between FDI and inequality in Latin America.

[Table 1 about here]

4.1.2. Panel estimates of the long-run FDI-inequality coefficient

We follow MacDonald and Ricci (2007) and Nowak et al. (2012) by implementing a dynamic ordinary least squares (DOLS) procedure to identify the long-run relationship between FDI and inequality. Kao and Chiang (2000) have shown that the panel version of the DOLS time-series estimator is less biased than other panel cointegration estimators, such as the panel version of the fully modified OLS (FMOLS) estimator. The panel DOLS estimator we use has the following form:

$$EHII_{it} = a_{1i} + a_2 \left(\frac{FDI}{GDP} \right)_{it} + \sum_{j=-p}^p \Phi_j \Delta \left(\frac{FDI}{GDP} \right)_{it-j} + \varepsilon_{it} \quad (2)$$

where Φ_j are coefficients of lead and lag differences which account for possible serial correlation and endogeneity of the regressors, thus yielding unbiased estimates. We estimate Equation (2) with fixed effects and fixed effects plus time dummies (to control for common time effects).¹⁹ The results are reported in the first two columns of Table 2 (where, for brevity, we show only the estimated slope coefficients).

[Table 2 about here]

¹⁹ We also included some impulse dummies to achieve a normal distribution of the residuals.

The coefficients are significant at the five percent level or better. On average, a percentage point increase in the FDI-to-GDP ratio increases inequality in terms of the Gini index by roughly 0.12 units when omitting time effects in the first column. The results of the model with country and time fixed effects indicate an impact that is somewhat lower, but still large. The panel cointegration results thus support the reasoning of Feenstra and Hanson (1997), who argue that inward FDI increases the relative demand for skilled labor in developing host countries. Higher relative wages, in turn, lead to increasing income inequality among households.

However, within-dimension based estimators may produce inconsistent and misleading results when the true slope coefficients are heterogeneous, as discussed in Section 3. To allow the slope coefficients to vary across countries, we use the between-dimension, group-mean panel DOLS estimator suggested by Pedroni (2001). This estimator involves estimating separate DOLS regressions for each country and averaging the long-run coefficients $\hat{a} = N^{-1} \sum_{i=1}^N \hat{a}_i$. The t -statistic for the average is the sum of the individual t -statistics divided by the root of the number of cross-sectional units, $t_{\hat{a}} = \sum_{i=1}^N t_{\hat{a}_i} / \sqrt{N}$.

The result can be found in the third column of Table 2. Within our Latin American sample, an increase in the FDI-to-GDP ratio by one percentage point increases the Gini index by 0.055 units. The magnitude of the estimated long-run coefficient is smaller than the within-dimension based panel coefficients, but the impact is still significant at the one percent level. In the fourth column, we account for common time effects using cross-sectionally demeaned data (by subtracting cross-sectional means from the observed data). This is equivalent to using the residuals from regressions of each variable on time dummies in place of the original variables. As can be seen, the impact is quantitatively smaller in the fourth column of Table 2, compared to the

second column. Once again, however, the estimated FDI-inequality coefficient is highly significant. We thus conclude that the effect of FDI is robust to the choice of different estimators.

4.1.3. Causality

The positive coefficient on $(FDI/GDP)_{it}$ does not necessarily reflect a causal effect of FDI on inequality; causality may also run from EHI_{it} to $(FDI/GDP)_{it}$ when FDI is attracted by wage dispersion in the host economy. Larger income inequality, i.e. a higher Gini coefficient, may reflect a decline in the real wages of less skilled workers. Multinational enterprises may then undertake (vertical) FDI and locate their low skilled activities in countries with a higher level of inequality in order to take advantage of lower wages for less skilled workers.

To test for the direction of causality, we include the (lagged) residuals,

$$ec_{it} = EHI_{it} - \hat{a}_{1i} - \hat{a}_2 \left(\frac{FDI}{GDP} \right)_{it}, \quad (3)$$

from DOLS long-run relationships (in Table 2) as error-correction terms into a vector error correction model (VECM) (estimated with one lag) of the form

$$\begin{bmatrix} \Delta EHI_{it} \\ \Delta \left(\frac{FDI}{GDP} \right)_{it} \end{bmatrix} = \begin{bmatrix} c_{1i} \\ c_{2i} \end{bmatrix} + \sum_{j=1}^p \Gamma_j \begin{bmatrix} \Delta EHI_{it-j} \\ \Delta \left(\frac{FDI}{GDP} \right)_{it-j} \end{bmatrix} + \begin{bmatrix} \alpha_1 \\ \alpha_2 \end{bmatrix} ec_{it-1} + \begin{bmatrix} \varepsilon_{1it} \\ \varepsilon_{2it} \end{bmatrix}, \quad (4)$$

where the c_i s are fixed effects; the error-correction term (ECT), ec_{it-1} , represents the error in, or deviation from, the equilibrium; and the adjustment coefficients α_1 and α_2 capture how EHI_{it} or $(FDI/GDP)_{it}$ respond to deviations from the equilibrium relationship. From the Granger representation theorem, we know that at least one of the adjustment coefficients must be non-zero if a long-run relationship between the variables is to hold. A significant adjustment coefficient also implies long-run Granger causality and thus long-run endogeneity (Hall and Milne 1994),

whereas a non-significant adjustment coefficient implies long-run Granger non-causality from the independent to the dependent variable(s), as well as weak exogeneity.

The front column of Table 3 indicates the panel estimation procedure on which the calculation of the ECTs is based. The subsequent columns show the t -statistics of the ECT with either inequality or FDI as the dependent variable. The results clearly indicate that causality runs from FDI to inequality. More specifically, the long-run causality appears to be unidirectional, implying that increased income inequality among households is the consequence and not the cause of inward FDI.

[Table 3 about here]

4.2. Panel cointegration results: extended sample and period of observation

By relying on EHII data, our results reported so far covered the period 1980-2000 and a small Latin American sample of five host countries for which a balanced dataset is available from this preferred source. While the data from this source can be regarded as most reliable, the small sample and short period of observation may render it impossible to generalize our baseline results for Latin America as a whole. To overcome this limitation, we draw on alternative sources of data on income inequality in the following.

Specifically, we expand country coverage and extend the period of observation by using the Gini coefficient (based on net income) from the SWIID (2011) developed by Solt (2009). As pointed out by Herzer and Nunnenkamp (2012), the SWIID combines information from the World Income Inequality Database (WIID) provided by the World Bank with information from the Luxembourg Income Study (LIS) database, which offers harmonized micro-data collected from multiple countries, and data from UNU-WIDER to create a dataset with greater coverage

than the LIS data and greater comparability than the WIID. The data provided by Solt (2009) are estimated (like the EHII data), and missing values are imputed to obtain greater coverage. Imputations rely on predicted adjustment factors to avoid missing observations.

Making use of the SWIID allows us to extend the cross-section as well as the time dimension of our investigation. We continue to rely on cointegration techniques in this section. Consequently, the time span now reaches from 1980 to 2006, i.e., the period that maximizes the observations under the constraint to provide a balanced panel.²⁰ At the same time, we now cover eleven countries in the region, including the largest economies in terms of GDP and population: Argentina, Brazil, Chile, Colombia, Costa Rica, Guatemala, Mexico, Panama, El Salvador, Uruguay, and Venezuela.

Given the heterogeneity of this larger sample and against the backdrop that the Latin American region experienced economic and financial turmoil in the 1980s, we use the group-mean panel estimator from Pedroni (2001), which allows the slope coefficients to vary across countries. We estimate Equation (2) with fixed effects and fixed effects plus time dummies.

[Table 4 about here]

The results can be found in Table 4. Both coefficients are significant at the five percent level. Importantly, the magnitude of the coefficients resembles the corresponding results for the smaller EHII sample (Table 2). An increase in the FDI-to-GDP ratio by one percentage point is associated with an increase of the Gini by 0.06 to 0.07 units. Hence, the inequality-increasing effect of FDI is somewhat stronger for the larger sample and the extended period of observation, compared to our earlier findings based on the EHII data.

²⁰ See below in Section 4.3. for a further extension of the period of observation up to 2011, for an unbalanced panel of Latin American host countries.

4.3. Unbalanced regression results

As noted above, our preferred panel cointegration approach addresses several estimation problems that often plague empirical work, including omitted variables, endogeneity and measurement error. The drawback is that this approach requires a balanced panel. This constrains both the sample of Latin American host countries and the period of observation, even when using the SWIID as in the previous section. Therefore, we employ an alternative estimation strategy in this section to test for the robustness of our major results.

Specifically, we follow Choi (2006) by estimating the following relationship for an unbalanced panel:

$$\begin{aligned} Gini_{it} = & \beta_0 + \beta_1 \ln \frac{FDI}{GDP_{it}} + \beta_2 \ln GDP_{it} + \beta_3 \ln GDPpc_{it} + \beta_4 \ln GDP_{it}^2 \\ & + \beta_5 Growth_{it} + \beta_6 \ln School_{it} + \gamma_i + \lambda_t + \vartheta_{it} \end{aligned} \quad (5)$$

where $Gini_{it}$ is the Gini coefficient for country i in year t . As in the previous section, we use the inequality measures from SWIID in order to include as many Latin American host countries as possible and cover the most recent past. FDI-to-GDP-ratios are from UNCTAD as before. In line with Choi (2006), we consider a number of control variables by including the size of the host country (GDP_{it}), its average income ($GDPpc_{it}$), and the growth rate of GDP per capita ($Growth_{it}$). We also control for human capital formation by including secondary gross school enrollment ($School_{it}$). Data on all these variables come from the World Development Indicators.²¹ Except for the growth rate of GDP per capita, all other control variables are included in natural logarithms.

This estimation approach allows us to broaden country coverage to an unbalanced panel of 23 countries in Latin America for the period from 1980 to 2011. The sample covers essentially the whole region and includes: Argentina, Belize, Bolivia, Brazil, Chile, Colombia, Costa Rica,

²¹ <http://data.worldbank.org/data-catalog/world-development-indicators> (downloaded: November 2013)

Cuba, Dominican Republic, Ecuador, El Salvador, Guatemala, Guyana, Haiti, Honduras, Mexico, Nicaragua, Panama, Paraguay, Peru, Suriname, Uruguay, and Venezuela.

[Table 5 about here]

Table 5 reports the results. Following Choi (2006), we always control for common time effects, by including dummy variables for each year. Every other column reports the results when additionally accounting for host-country fixed effects. The evidence on GDP and GDP per capita resembles the findings of Choi (2006) without host-country fixed effects in columns (1) and (3), while the coefficients of these two control variables are no longer significant at conventional levels when accounting for host-country fixed effects in columns (2) and (4). The growth rate of GDP per capita enters insignificant throughout.

Importantly, our variable of principal interest, FDI, is not affected by modifications in the specification of the unbalanced panel regression. The coefficient on FDI proves to be significantly positive in all estimations, at the ten percent level or better. In other words, Table 5 corroborates the inequality-increasing effects of FDI, independent of whether we exclude or include host-country fixed effects in columns (1)-(4). This effect also carries over to columns (5)-(6) where we augment the specification by $School_{it}$ to account for human capital formation. The last two columns report the results when we include GDP per capita in quadratic form to allow for a potential Kuznets-style relationship. Again, the results on FDI are significant and the coefficients are positive, revealing an inequality-increasing effect in a reasonable range.

4.4. Results for the recent past

In the final step of our empirical analysis, we replicate the estimations reported for the extended samples in the two previous sections by focusing on the more recent past and excluding the

1980s. The 1980s are often referred to as the ‘lost decade’ in large parts of Latin America. Political crises and financial turmoil were associated with poor economic performance and may have affected the link between FDI and income inequality, possibly resulting in biased estimation results. Furthermore, the stability of the FDI-inequality link over a fairly long period of time is open to debate in the light of the recent trend of declining inequality in various Latin American host countries (López-Calva and Lustig 2010; Gasparini and Lustig 2011). We assess this issue at least tentatively by excluding the distant past from the estimations.

[Table 6 about here]

For a start, we modify the panel cointegration analysis reported in Section 4.2 by gradually excluding two years of the 1980s from the estimation. As can be seen in Table 6, this hardly affects our major result on the inequality-increasing effects of FDI in the first three steps of excluding the most distant past. However, our panel cointegration results for the sample of eleven Latin American countries prove to be sensitive to the exclusion of the late 1980s. The effect of FDI on income inequality loses its significance completely when the period of observation is restricted to 1990-2006. Unless more data become available for a balanced panel of Latin American host countries, it remains open to question whether this finding is only because of the relatively short time dimension that remains after excluding the 1980s, or whether the last columns in Table 6 reveal a change over time in the FDI-inequality link.

[Table 7 about here]

The modified regression results for the unbalanced panel of 23 Latin American countries including the most recent past speak against a pronounced change over time in the FDI-inequality link.²² To the contrary, most of the coefficients on FDI reported in Table 7 are slightly higher and

²² For the sake of brevity, we show only the results after excluding the 1980s in one single step.

their significance levels increased, compared to the corresponding estimations shown in Table 5 above. In other words, the exclusion of the 1980s from the unbalanced panel regressions mitigates concerns that our previous results covering the whole period 1980-2011 were distorted by the crises during the ‘lost decade.’

5. Conclusion

We analyzed whether foreign direct investment (FDI) has contributed to the wide income gaps in Latin America. We applied panel cointegration techniques to assess the long-run impact of inward FDI stocks on income inequality among households in different samples of Latin American host countries since the early 1980s. We also employed unbalanced panel regressions to cover essentially the whole region as well as the most recent past. The panel cointegration analysis revealed a significant and positive effect on income inequality. Our results proved to be robust to the choice of different estimation methods, and the unbalanced regression analysis underscored our major finding. We did not find evidence for reverse causality from inequality to FDI.

These findings suggest that the North-South model of Feenstra and Hanson (1997) does not only hold for Mexico and the free trade conditions prevailing among NAFTA members. The model’s predictions also hold more generally for different samples of Latin American countries, including host countries where trade liberalization proceeded sooner (e.g., Chile) or later (e.g., Argentina, Bolivia, and Brazil) on a unilateral basis. According to the Feenstra-Hanson model, FDI increases the demand for relatively skilled labor in developing countries hosting FDI from more advanced source countries.

Our major result on the inequality-increasing effects of FDI involves a policy dilemma for Latin American governments. Many governments in the region promote FDI inflows to benefit

from spillovers and stimulate economic growth. At the same time, the alleviation of poverty and income inequality has figured high on the policy agenda of Latin American governments since recently. This raises the question of how these potentially conflicting goals may be reconciled. One option to address the policy dilemma might be to rely on strong economic growth, which appears to have contributed to the recent decline in inequality in Latin America (Gasparini and Lustig 2011). However, FDI-induced growth may be insufficient to contain social conflicts as long as FDI benefits primarily relatively skilled workers.

Another option might be to target FDI inflows, e.g., by selective approvals, in order to avoid higher FDI-induced demand for skilled workers. Indeed, UNCTAD (2013: 1) observes a trend toward more restrictive FDI regulations “continuing for many years now.” All the same, it appears unlikely that this trend will shift the composition of FDI in favor of relatively unskilled-labor intensive operations in Latin American host countries. Targeting FDI in unskilled-labor intensive industries is rather illusory as long as policymakers are mainly interested in technology spillovers in export-oriented and high-tech industries where the demand for higher skills is particularly strong.

More realistically, the recent wave of social policies in large parts of the region may help address the above noted policy dilemma by containing social conflicts and improving basic education. As argued by Birdsall et al. (2011), progressive social transfers including targeted spending benefited a large proportion of households at the bottom of the income distribution. Some of these innovative schemes made transfers conditional on children’s school attendance. In this way, they can contribute to the decline in the premium to skills by expanding the pool of workers with at least basic education.

Better education at all levels offers a possible win-win strategy. It would not only improve the supply of skilled labor, but also help attract higher FDI. Education has traditionally been

neglected in many Latin American countries, and an insufficiently qualified workforce appears to represent a major bottleneck to higher FDI flows to the region (Donaubauer et al. 2013). Hence, more and better schooling and improving the qualification of the workforce should figure high on the policy agenda, in order to narrow the gap between the demand and supply of sufficiently skilled labor. This, in turn, could allow for a smoother “transition to a new technological paradigm” (Aghion and Howitt 1998).

It remains to be seen whether recent trends of increasing access to education and diminishing income gaps will prove persistent and common throughout Latin America. While our analysis provided few indications that the FDI-inequality link has changed already, it clearly deserves attention in future research to re-assess this possibility. Longer time series would be required to evaluate whether FDI-induced income inequality is a transitional phenomenon in Latin America, which recedes to the extent that host countries successfully manage an FDI-induced technological transition and strengthen the supply of sufficiently skilled labor.

Other issues for further research include the question of whether the distributional effects depend on the structure, type, and origin of inward FDI. For instance, it could provide further insights if inward FDI is differentiated by sectors and industries. While theoretical models tend to focus on wage disparity in the manufacturing sector, FDI in the services sector has played an increasingly important role, including in developing host countries. Furthermore, South-South FDI has increasingly supplemented North-South FDI, and the distributional consequences of the former may differ from the latter.

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Table 1 Panel cointegration tests

Pedroni (1999, 2004)	
Panel PP t -statistic	-4.87***
Panel ADF t -statistic	-4.87***
Group PP t -statistic	-5.89***
Group ADF t -statistic	-5.13***
Kao (1999)	
DF_{ρ} statistic	-8.15***
DF_t statistic	-5.22***
ADF t -statistic	-3.98***
DF_{ρ}^* statistic	-3.53***
DF_t^* statistic	-3.89***

Notes: *** indicate a rejection of the null of no cointegration at the 1% significance level. The number of lags is based on the Schwarz information criterion with a maximum number of four.

Table 2 Estimates of the long-run effects of FDI on inequality

Within-dimension DOLS estimator		DOLS mean group estimator	
Kao and Chiang (2000)		Pedroni (2001)	
Without period effects	Period effects	Without period effects	Period effects
0.123***	0.083**	0.055***	0.033***
(3.83)	(2.42)	(7.54)	(2.63)

Notes: **(***) indicate a rejection of the null of no cointegration at the 5% (1%) significance level.

Table 3 VECM: Long-run causality test in the panel

	Dependent variable ΔEHI_{it}	Dependent variable $\Delta \left(\frac{FDI}{GDP} \right)_{it}$
Calculation method of ECT		
Kao and Chiang (2000)		
Fixed effects	-3.99***	1.01
Two-way fixed effects	-3.89***	1.05
Pedroni (2001)		
Fixed effects	-5.05***	-1.13
Two-way fixed effects	-6.07***	1.24

Notes: *** indicate a rejection of the null of no cointegration at the 1% significance level.

Table 4 Estimates of the long-run effects of FDI on inequality using the SWIID data

DOLS mean group estimator	
Pedroni (2001)	
Without period effects	Period effects
0.072**	0.059**
(2.02)	(1.99)

Notes: ** indicate a significance at the 5% level.

Table 5 Unbalanced panel results (1980 -2011)

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
FDI	0.226* (0.127)	0.156* (0.0944)	0.249* (0.130)	0.208** (0.0953)	0.373* (0.198)	0.333** (0.133)	0.385* (0.198)	0.388*** (0.135)
GDP	1.090*** (0.130)	0.0658 (2.346)	1.055*** (0.132)	-0.867 (2.392)	1.195*** (0.176)	-4.750* (2.690)	1.242*** (0.180)	-4.554* (2.674)
GDP p.c.	-5.169*** (0.350)	-1.944 (2.289)	-5.172*** (0.357)	-1.007 (2.317)	-6.646*** (0.538)	1.933 (2.619)	4.920 (9.710)	-34.98** (16.69)
GDP p.c. squared							-0.732 (0.614)	2.213** (0.989)
GDP p.c. growth			6.242 (4.531)	-2.720 (2.654)	5.518 (5.512)	-1.196 (3.122)	5.284 (5.511)	-0.263 (3.130)
Schooling					1.380 (0.846)	3.309*** (1.074)	1.353 (0.845)	4.166*** (1.133)
Constant	60.98*** (2.818)	60.67 (40.23)	62.92*** (2.811)	76.00* (41.13)	65.57*** (4.014)	131.9*** (45.29)	18.99 (39.25)	276.8*** (78.83)
Fixed Effects	No	Yes	No	Yes	No	Yes	No	Yes
Period Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	507	507	495	495	363	363	363	363
R-squared	0.364	0.151	0.363	0.165	0.391	0.239	0.394	0.251

Notes: Standard errors are reported in parentheses. ***, **, * indicate a significance to the 1%, 5% and 10% level.

Table 6 DOLS mean group estimates of the long-run effects of FDI on inequality

Period of observation	Fixed	Fixed/Period
1982-2006	0.072*** (4.13)	0.059* (1.87)
1984-2006	0.061*** (3.79)	0.072** (2.39)
1986-2006	0.047*** (2.67)	0.064* (1.75)
1988-2006	0.034** (2.16)	0.017 (0.42)
1990-2006	0.0250 (1.55)	-0.002 (-0.07)

Notes: *, **, *** indicate a significance at the 10%, 5% and 1% level.

Table 7 Unbalanced panel results (1990 -2011)

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
FDI	0.285** (0.129)	0.190** (0.0965)	0.279** (0.129)	0.191** (0.0964)	0.450** (0.205)	0.341*** (0.127)	0.488** (0.208)	0.376*** (0.126)
GDP	0.820*** (0.132)	2.962 (3.225)	0.830*** (0.134)	3.148 (3.231)	0.996*** (0.184)	-5.021 (3.701)	1.031*** (0.187)	-5.059 (3.662)
GDP p.c.	-4.505*** (0.339)	-7.694** (3.193)	-4.537*** (0.350)	-7.995** (3.189)	-5.653*** (0.530)	-3.047 (3.717)	4.431 (9.591)	-40.48** (16.04)
GDP p.c. squared							-0.641 (0.608)	2.271** (0.947)
GDP p.c. growth			3.056 (5.196)	-3.820 (2.905)	4.371 (6.575)	0.471 (3.320)	4.006 (6.582)	1.370 (3.306)
Schooling					0.832 (0.935)	3.965*** (1.127)	0.889 (0.936)	4.728*** (1.160)
Constant	61.08*** (2.689)	35.42 (54.92)	61.12*** (2.709)	33.31 (55.29)	60.89*** (4.285)	172.6*** (60.97)	20.03 (39.04)	323.8*** (87.25)
Fixed Effects	No	Yes	No	Yes	No	Yes	No	Yes
Period Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	372	372	371	371	267	267	267	267
R-squared	0.398	0.233	0.396	0.241	0.412	0.373	0.415	0.389

Notes: Standard errors are reported in parentheses. ***, **, * indicate a significance to the 1%, 5% and 10% level.

Table A.1 Summary Statistics

Country	Average of $\left(\frac{FDI}{GDP}\right)_i$	Average of $EHII_i$	Average of $GDP\ p.c.$
<i>Bolivia</i>	23.9	47.1	934.3
<i>Chile</i>	48.0	45.7	3363.8
<i>Colombia</i>	7.1	43.1	2293.0
<i>Mexico</i>	8.4	41.3	5067.0
<i>Uruguay</i>	6.4	42.8	5708.1

Table A.2 Augmented Dickey-Fuller Test Results

Levels		
Deterministic terms: <i>constant, trend</i>		
Variables	Country	<i>t</i> -stat
<i>Inequality</i>	<i>Bolivia</i>	-2.44
	<i>Chile</i>	-2.42
	<i>Colombia</i>	-2.66
	<i>Mexico</i>	-2.15
	<i>Uruguay</i>	-1.74
<i>FDI</i>	<i>Bolivia</i>	0.61
	<i>Chile</i>	-1.82
	<i>Colombia</i>	-3.16
	<i>Mexico</i>	-2.18
	<i>Uruguay</i>	-2.50
<i>Critical Values 5%(10%)</i>		-3.69(-3.29)

Notes: Critical values are from MacKinnon (1996). The single-country ADF-test equations allow for two lagged differences of the endogeneous variable to correct for potential autocorrelation in the residuals.