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FDI and Income Inequality - Evidence from Latin American Economies

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FDI and Income Inequality - Evidence from Latin American Economies

Dierk Herzer, Philipp Hühne and Peter Nunnenkamp

Abstract:

We analyze whether foreign direct investment (FDI) has contributed to the typically wide income gaps in five Latin American host countries. We perform country-specific and panel cointegration techniques to assess the long-run impact of inward FDI stocks on income inequality among households in Bolivia, Chile, Colombia, Mexico and Uruguay. The panel cointegration analysis reveals a significant and positive effect on income inequality. Furthermore, FDI contributed to widening income gaps in all individual sample countries, except for Uruguay. Our findings are robust to the choice of different estimation methods. There is no evidence for reverse causality.

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 country-specific and panel cointegration analyses to assess the distributional effects of inward FDI in five Latin American countries – Bolivia, Chile, Colombia, Mexico, and Uruguay – during the 1980-2000 period. 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

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.

literature reveals that “almost all evidence shows that FDI and foreign ownership are associated 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

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. See also Herzer and

countries lacking HQ services of domestic MNEs tend to be at the other end of the spectrum of 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.

Nunnenkamp (2011), who find that FDI in advanced European host countries reduced income inequality in the long run.

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).

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 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 (Lipsev 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

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.

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”, 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 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. 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.

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 later reducing inequality.¹¹ 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

¹¹ 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.

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, Choi (2006) finds more pronounced income inequality where the ratio of FDI stocks to GDP is higher. The estimations of Basu and Guariglia (2007) for a 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. 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 cointegration techniques both in a panel context and for individual countries. 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:¹²

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

where $EHII_{it}$ 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.

¹² When not further specified, the term inequality refers to income inequality among households.

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. We also run country-specific regressions to examine the impact of FDI on inequality for each country individually.

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>). 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).

Our analysis covers the period from 1980 to 2000 (21 yearly observations per country). This is the longest time span available to conduct an empirical analysis with a balanced panel. We include all Latin American countries with complete time series data over this period: Bolivia, Chile, Colombia, Mexico and Uruguay.

In our view, our time series are sufficiently long to conduct a cointegration analysis. Several cointegration analyses for individual countries are based on shorter periods (e.g., Crombrughe et al. 1997; Irvin and Izurieta 2000). However, it should be mentioned that the behavior of the individual country test statistics we use (in Section 4.2) may be affected by the small sample size. To deal with this problem, we use finite-sample critical values. In addition, we use several test and estimation methods to ensure the robustness of our results. Specifically, we use panel cointegration methods (in Section 4.1), 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.

[Table 1 about here]

Table 1 provides summary statistics for the five Latin American economies in our sample. We also add averages of per capita GDP from the World Development Indicators (WDI; <http://data.worldbank.org>) to give an impression of the state of development of the particular

economy.¹³ 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 (see also the Introduction). Chile and Bolivia are the top FDI recipients. The FDI-to-GDP ratio in Chile is more than seven times higher than the corresponding ratio in Uruguay, which represents the tailight in terms of FDI in our sample.

As discussed 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 the Appendix.

4. Empirical analysis

We first use panel cointegration techniques to examine the “average” relationship between FDI and inequality for the five Latin American countries. Then, we employ time series cointegration methods to investigate the FDI-inequality relationship for each of the five Latin American countries individually.

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

¹³ Per capita GDP is in prices of 2000; this information does not enter our empirical analysis, however.

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{e}_{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 2. 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 2 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).¹⁴ The results are reported in the first two columns of Table 3 (where, for brevity, we show only the estimated slope coefficients).

[Table 3 about here]

¹⁴ We also included some impulse dummies to achieve a normal distribution of the residuals; see Section 4.2.1 for details.

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 effect 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, $t_{\hat{a}} = \sum_{i=1}^N t_{\hat{a}_i} / \sqrt{N}$.

[Table 3 about here]

The result can be found in the third column of Table 3. 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 3, compared to the

second column. Once again, however, the estimated FDI-inequality coefficient is highly significant. We thus conclude that the effect of FDI on 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,

$$ec_{it} = EHI_{it} - \hat{\alpha}_1 - \hat{\alpha}_2 \left(\frac{FDI}{GDP} \right)_{it}, \quad (3)$$

from DOLS long-run relationships (in Table 3) 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 4 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 4 about here]

4.2. Individual country results

The results from the panel estimations reported so far suggest that FDI increases inequality on average. However, this effect may vary among the countries in our sample. As in the previous section, we start with testing for cointegration. We then apply several techniques to estimate the long-run impact of FDI on inequality in each country in our sample. We also test the robustness of our results and the direction of causality.

4.2.1. Individual time series cointegration tests

We rely on the two most common approaches to test for cointegration: The single-equation two-step procedure proposed by Engle and Granger (1987), and the single-equation (conditional) error correction model (ECM) test procedure based on the work of Ericsson and MacKinnon (2002).

As is well known, the approach of Engle and Granger (1987) involves running the static cointegrating relationship given by Equation (1) (for each country) and testing the residuals $\hat{\epsilon}_t$

for stationarity using a standard ADF-regression (without intercept). The lag length in the ADF-test is determined using the t -sig method, i.e. downward testing an arbitrarily large number of lags, in our case four. The results of the cointegration tests are reported in Table 5. To account for the small sample size, we calculated the critical values for the Engle-Granger cointegration test using the (“small sample”) response surface estimates from MacKinnon (2010). As can be seen, the null of no cointegration is rejected at conventional significance levels. This is true for four of our five countries. Uruguay seems to be an exception.

[Table 5 about here]

A well-known problem with the Engle-Granger procedure is that it imposes a common factor restriction by restricting the long-run elasticities to be equal to the short-run elasticities. If this restriction is invalid, residual-based cointegration tests may suffer from low power. Therefore, we also use the standard ECM cointegration test, which allows the long-run elasticities to differ from the short-run elasticities and hence does not impose a possibly invalid common factor restriction. Specifically, we estimate a conditional error-correction model of the form

$$\Delta EHH_t = b_1 + b_2 EHH_{t-1} + b_3 \left(\frac{FDI}{GDP} \right)_{t-1} + \sum_{i=1}^p \eta_i \Delta EHH_{t-i} + \sum_{i=0}^p \gamma_i \Delta \left(\frac{FDI}{GDP} \right)_{t-i} + u_t \quad (5)$$

Following common practice, we eliminate the insignificant short-run dynamics (lagged differences) successively according to their lowest p -values. A significantly negative coefficient of the lagged dependent level variable, b_2 , indicates cointegration. Accordingly, the null of no cointegration to be tested is $b_2 = 0$. The corresponding finite sample critical values can be calculated from the (“small sample”) response surfaces in Ericsson and MacKinnon (2002).

Table 6 reports our results, where, for brevity, we report only the t -statistics of the error correction coefficients. As noted in the third column of Table 6, we included impulse dummies,

y_{XX} , whenever necessary, to account for large outliers in the residuals and, thereby, achieve a normal distribution of the residuals; y_{XX} is one in the year 19XX and is zero otherwise.

[Table 6 about here]

The ECM cointegration test is in line with the Engle-Granger approach. Therefore, we can safely conclude that there is a cointegrating relationship between FDI and inequality in Bolivia, Chile, Colombia, and Mexico, while there is no cointegrating relationship in the case of Uruguay.

Uruguay's low FDI-to-GDP ratio (Table 1) may provide a possible explanation for the lack of cointegration.¹⁵ Traditionally, FDI played a marginal role in Uruguay and may have been too small to have an impact (del Castillo and García 2012). Moreover, Mercosur partner countries, notably Argentina, contributed a considerable share of FDI in neighboring Uruguay.¹⁶ FDI from these sources does not fit into the Feenstra-Hanson framework and could have weakened the relation between FDI and inequality.

4.2.2. Individual country estimates of the long-run FDI-inequality coefficient

In order to estimate the long-run impact of FDI on inequality, we use the conventional time series DOLS estimator of Stock and Watson (1993), which is known to perform well in small samples. The good small sample properties are particularly useful in our case because of the limited data available. The DOLS time series procedure involves estimating Equation (2) for each country separately.

We report the estimated long-run FDI-inequality coefficients for each country in Table 7 along with their t -statistics and included impulse dummies (to achieve a normal distribution of

¹⁵ Lacking evidence for cointegration notwithstanding, we estimate the long-run impact of FDI on inequality for Uruguay, too. However, we do not expect a statistically significant coefficient in the case of Uruguay because of the absence of a cointegrating relationship.

¹⁶ For details, see <http://www.uruguayxxi.gub.uy/wp-content/uploads/2012/04/FDI-in-Uruguay-April-2012-URUGUAY-XXI-.pdf> (accessed: August 2012).

the residuals). In order to make sure that our DOLS models are well specified, we also present several diagnostic test statistics. All these test statistics, except those for Uruguay, indicate that the estimated models are statistically well-specified: the Lagrange multiplier (*LM*) tests for autocorrelation based on 1 and 3 lags, respectively, do not indicate any problems concerning autocorrelated residuals; the models also pass the ARCH tests of autoregressive conditional heteroscedasticity of order $k = 1, 3$; and the Jarque Bera tests (*JB*) cannot reject the hypothesis of normally distributed residuals. The fact that Uruguay does not pass the diagnostic tests is consistent with our finding that there is no cointegrating relationship between FDI and inequality in Uruguay.

[Table 7 about here]

The estimates suggest that increased FDI typically leads to increased income inequality among households in the Latin American host countries. In particular, our country-specific findings are in line with those of Nunnenkamp et al. (2007) for Bolivia as well as Aitken et al. (1996), Feenstra and Hanson (1997), and Hanson (2003) for Mexico. According to Table 7, inequality in Bolivia appears to react most strongly to the presence of foreign enterprises. On average, the Gini index in Bolivia increases by 0.394 units when the FDI-to-GDP-ratio rises by one percentage point. The long-run coefficient is slightly smaller for Mexico. The impact of FDI appears to be considerably weaker in Chile and Colombia, though still significantly positive at the one percent level. It should be taken into account, however, that the average FDI-to-GDP ratio is highest in Chile, while the ratio in Colombia is almost as low as in Uruguay (Table 1). This implies that the increase in income inequality resulting from an increase in the FDI-to-GDP ratio by ten percent is considerably stronger in Chile (0.55 units of the Gini index) than in

Colombia (0.10 units).¹⁷ Apart from the low FDI-to-GDP ratio, the comparatively weak impact in Colombia may be attributed to the country's limited openness to trade during the period of observation.¹⁸

4.2.3. Robustness and causality

We now examine the robustness of our results to different estimation techniques. Table 8 reports the results of the conventional time series FMOLS estimator of Phillips and Hansen (1990). In addition, we present the long-run coefficients from the ECM procedure. These were calculated from the individually estimated ECM as $a_{2i} = b_{3i} / |b_{2i}|$. The FMOLS results are largely in line with the results in the previous section.¹⁹ The long-run coefficients have the same magnitude as our DOLS coefficients for Chile, Colombia, and Mexico.²⁰ This is also true for the long-run coefficients from the ECM procedure.

[Table 8 about here]

Finally, we investigate the direction of causality for each individual country. The causality testing procedure is analogous to that discussed in the previous section: We enter the (lagged) residuals from our individual country DOLS regressions as error correction terms into a VECM. As before, we use one lag given the short sample period.

[Table 9 about here]

¹⁷ The increase in income inequality resulting from an increase in the FDI-to-GDP ratio by ten percent in Chile is even stronger than the corresponding calculation for Mexico (0.31 units).

¹⁸ As mentioned in Section 2 above, Das (2002) argues that the predictions of the model of Feenstra and Hanson (1997) critically depend on openness to trade. It fits into this reasoning that Colombia reported the lowest average import-to-GDP ratio among the five sample countries during the period of observation (for details, see the World Development Indicators). Furthermore, Colombia was characterized by comparatively high average tariffs and hidden import barriers as well as shortage of foreign exchange for importing at the official exchange rate (World Economic Forum 2000).

¹⁹ The exception is Bolivia where the coefficient on FDI loses its significance in Table 8.

²⁰ Once again, the long-term coefficient does not reveal a significant impact of FDI in Uruguay.

The error correction terms in the first column of Table 9 with income inequality as the dependent variable are highly significant (again with the exception of Uruguay). By contrast, we do not find causality running in the opposite direction. The error correction terms in the equation with $\Delta(FDI / GDP)_{it}$ as the dependent variable are insignificant for all countries. This is in line with the findings from our panel analysis: FDI seems to cause changes in income distribution, while there is no evidence of reverse causation.

5. Conclusion

We analyzed whether foreign direct investment (FDI) has contributed to the wide income gaps in five Latin American host countries. We applied country-specific and panel cointegration techniques to assess the long-run impact of inward FDI stocks on income inequality among households in Bolivia, Chile, Colombia, Mexico and Uruguay during the period from 1980 to 2000. The panel cointegration analysis revealed a significant and positive effect on income inequality. Furthermore, FDI contributed to widening income gaps in all individual sample countries, except for Uruguay. Our results proved to be robust to the choice of different estimation methods. 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 for Latin American countries where FDI contributed considerably to the host countries' production capacity and where trade liberalization proceeded sooner (Chile) or later (Bolivia) 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. This implies a major policy challenge for many Latin American host countries where education has been neglected and skilled labor is in short supply. 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).

At the same time, our analysis provides some indications that the distributional effects depend on the magnitude, structure and type of inward FDI. Most obviously perhaps, it appears that FDI must exceed a certain threshold, in terms of its contribution to production capacity in the host country, to trigger significant distributional effects. The cases of Uruguay and Colombia suggest that inequality may be less affected when FDI comes from regional sources and host countries are less open to trade.

The role of the origin and type of FDI in shaping the distributional effects of FDI clearly deserves more research. This applies in particular to horizontal FDI in less open developing host countries. In the Latin American context, Argentina and Brazil would be interesting cases in point as they attracted considerable FDI inflows, even though they opened up to trade later than our sample countries. It could also provide further insights if inward FDI was 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. However, future research along these lines is impeded by persistent data limitations. The same is true for the assessment of changes over time in the distributional effects of FDI. In particular, 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.

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Table 1 Summary Statistics

Country	Average of $\left(\frac{FDI}{GDP}\right)_t$	Average of EHI_t	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 2 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 3 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 4 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 5 Engle Granger cointegration test results

Country	ADF-statistic
<i>Bolivia</i>	-3.82**
<i>Chile</i>	-5.84**
<i>Colombia</i>	-4.56**
<i>Mexico</i>	-4.62**
<i>Uruguay</i>	-0.13
<i>Critical</i>	-4.47 (-3.64)
<i>Values 1% (5%)</i>	

Notes: Critical values for the Engle-Granger cointegration test are from MacKinnon (2010).

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

Table 6 ECM cointegration test results

Country	ECM t -statistic	Dummy	\bar{R}^2
<i>Bolivia</i>	-6.27**	y90	0.83
<i>Chile</i>	-9.77**	y89, y94	0.95
<i>Colombia</i>	-6.14**	y91	0.82
<i>Mexico</i>	5.68**	y92	0.76
<i>Uruguay</i>	0.01	--	0.07
<i>Critical Values 1%(5%)</i>	-4.28(-3.41)		

Notes: Critical values for the cointegration test procedure are from Ericsson and MacKinnon (2002). ** indicate a rejection of the null of no cointegration at the 1% significance level.

TABLE 7 Dynamic OLS estimates

Country	Coefficient	t -statistic	Dummy	DIAGNOSTIC TESTS					
				\bar{R}^2	LM(1)	LM(3)	ARCH(1)	ARCH(3)	JB
<i>Bolivia</i>	0.394**	3.31	y90	0.87	0.14(0.72)	0.48(0.70)	3.02(0.10)	1.12(0.38)	0.53(0.77)
<i>Chile</i>	0.114**	6.82	y89, y94	0.97	0.05(0.82)	1.71(0.24)	3.18(0.09)	2.04(0.17)	0.83(0.66)
<i>Colombia</i>	0.144**	3.05	y91	0.81	1.20(0.30)	1.34(0.32)	0.07(0.80)	0.11(0.95)	0.10(0.95)
<i>Mexico</i>	0.364**	4.68	y92	0.74	0.06(0.81)	0.10(0.96)	0.03(0.87)	0.36(0.78)	3.27(0.19)
<i>Uruguay</i>	-0.740	-1.01	--	0.20	73.80(0.00)	22.21(0.00)	7.51(0.02)	10.54(0.00)	0.27(0.87)

Notes: ** indicate a rejection of the null of no cointegration at the 1% level. p -values are reported in parentheses

Table 8 Results from robustness checks

Country	Coefficient	FMOLS		ECM
		<i>t</i> -statistic	Dummy	long-run coefficient:
<i>Bolivia</i>	0.052	1.16	y90	0.068
<i>Chile</i>	0.112**	6.55	y89, y94	0.098
<i>Colombia</i>	0.136**	3.44	y91	0.111
<i>Mexico</i>	0.301**	4.18	y92	0.278
<i>Uruguay</i>	0.062	0.13	--	8.210

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

Table 9 Estimates of the long-run effects of FDI on inequality

Country	Dependent Variable	Dependent Variable
	ΔEHI_{it}	$\Delta \left(\frac{FDI}{GDP} \right)_{it}$
<i>ec_t Bolivia</i>	-4.26**	0.81
<i>ec_t Chile</i>	-4.97**	-0.11
<i>ec_t Colombia</i>	-4.34**	1.76
<i>ec_t Mexico</i>	-3.69**	1.70
<i>ec_t Uruguay</i>	-0.96	-0.50

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

Table A.1 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</i>		-3.69(-3.29)
<i>Values 5%(10%)</i>		

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