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The distributional effects of capital account liberalization

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ABSTRACT

Episodes of account liberalization increase the Gini measure of inequality, based on panel data estimates for 149 countries from 1970 to 2010. These episodes are also associated with a persistent increase in the share of income going to the top. We investigate three channels through which these impacts could occur. First, the impact of liberalization on inequality is stronger where credit markets lack depth and financial inclusion is low; positive impacts of liberalization on poverty rates also vanish when financial inclusion is low. Second, the impact on inequality is also stronger when liberalization is followed by a financial crisis. Third, liberalization seems to alter the relative bargaining power of firms and workers: the labor share of income falls in the aftermath of capital account liberalization.

1. Introduction

There has been an increase in global financial integration over the past fifty years, reflected for instance in a steady decline in the number of restrictions that countries impose on cross-border financial transactions. Indices of capital account openness show an increase, on average, across all income groups, with a particularly significant rise occurring at the beginning of the 1990s. The growth effects of this liberalization have been extensively studied but remain the subject of debate, with Henry (2007) claiming a positive impact on growth and Rodrik and Subramania (2009) taking a more skeptical view. This paper studies the distributional effects of capital account liberalization. While there is a vast literature on the distributional effects of trade (Helpman et al., 2015, and references cited therein), there are only a few studies that analyze the relation between financial globalization and inequality. This is surprising because, just as with trade, there are various channels through which capital account liberalization can affect inequality (Claessens and Perotti, 2007).

One channel is through the impact of liberalization on risk-sharing. In theory, financial openness should foster international risk-sharing and domestic consumption smoothing (Kose et al., 2009). In practice, the strength of financial institutions may play a crucial role in determining the extent to which this takes place. In countries with strong financial institutions, financial globalization may reduce inequality by allowing better consumption smoothing and lower volatility. But where financial institutions are weak and access to credit is not inclusive, liberalization may bias financial access in favor of those who are well off and therefore increase inequality.

A second channel is through the effect of liberalization on the likelihood of financial crises. On the one hand, financial crises may reduce inequality as bankruptcies and falling asset prices may have greater impact on those who are better off. On the other hand, financial crises associated with long-lasting recessions may disproportionately hurt the poor and hence increase inequality (de Haan and Sturm, 2016).[1]

Finally, capital account openness may affect the distribution of income through its effect on the bargaining power of labor. If capital account liberalization represents a credible threat to reallocate production abroad, it may lead to an increase in the profit-wage ratio and to a decrease in the labor share of income (Harrison, 2002).

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[1] The empirical evidence on the effect of financial crises on inequality is mixed. Baldacci et al. (2002) find that inequality increases in the aftermath of currency crises. Atkinson and Morelli (2011) find that inequality increased following the financial crisis in Sweden in 1991 and in Iceland in 2007, but decreased in the aftermath of the early 1990s crises in Norway and Finland. Agnello and Sousa (2012), using annual data for 62 economies over the period 1980–2006, find that inequality increases before banking crisis episodes and declines afterward. In contrast, de Haan and Sturm (2016), using panel as well as cross-country regressions, find that systemic banking crises are robustly associated with an increase in the Gini coefficient of about 1 percentage point.

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This paper contributes to the empirical literature linking finance and inequality.[2] The contributions of the paper are two-fold. First, we provide evidence that capital account liberalization raises inequality, using a large (unbalanced) panel dataset comprising 149 advanced, emerging and low-income economies from 1970 to 2010. The focus of much of the previous literature has been on within-country experience or on a more limited set of emerging market economies.[3] Second, we empirically examine the key mechanisms—such as the extent of financial development, the occurrence of financial crises, and the impact on labor's bargaining power—through which capital account liberalization may affect inequality. We show that each of these mechanisms is operative and needs to be considered in order to have a full picture of the distributional impacts of capital account liberalization.[4]

We check the robustness of our findings to the use of alternate measures of inequality. Specifically, we use the Gini coefficients, the income shares going to the top, poverty rates, and the labor share of income as alternate measures of the distribution of incomes. Given the weaknesses associated with any one measure of inequality, this is an important check on our findings. In addition, we conduct several robustness checks of our findings to address (i) omitted variable bias; (ii) endogeneity bias; and (iii) sensitivity to alternate econometric specifications. The effects of capital account liberalization may very well be confounded by other concurrent political and policy changes. One concern is that liberalization is often enacted by governments of the political right, who may simultaneously pursue other policies that tend to increase inequality. We show, however, that our results are robust to controlling for the political affiliation of governments. We also address endogeneity bias through the use of some novel instruments, including one that attempts to measure the peer pressure a country may feel to liberalize if its main trading partners are liberalizing. Our results are also shown to be robust to alternate econometric specifications. Our baseline results are based on the autoregressive distributed lag model, estimated both by OLS and GMM; in addition, we also use the local projection method of Jordà (2005).

The key findings of the paper are as follows. Episodes of capital account liberalization are associated with a statistically significant and persistent increase in inequality. In particular, we find that capital account liberalization has typically increased the Gini coefficient by about 0.8 percent in the very short term (1 year after the occurrence of a liberalization reform) and by about 1.4 percent in the medium term (5 years after). These episodes are also associated with a persistent increase in the share of income going to the top 1, 5 and 10 percent of the population.

The level of financial development and inclusion and the occurrence of crises play a key role in shaping the response of inequality to financial globalization. In particular, capital account liberalization leads to larger increases in inequality in countries when it is followed by financial crises or in countries with a weaker quality of financial institutions and low financial inclusion. We also find evidence that capital account liberalization lowers the labor share of income, consistent with the view that liberalization curtails the bargaining power of workers relative to firms. As noted earlier, these results are robust to different sets of controls, different estimation techniques, alternate measures of capital account openness and inequality, and checks for omitted variable and endogeneity biases.

The rest of the paper is organized as follows. The next section describes the data and descriptive statistics of the evolution of inequality and capital account openness. Section 3 analyzes the effect of capital account openness on inequality and provides some robustness checks. Section 4 empirically identifies some of the mechanisms through which capital account liberalization affects inequality. Section 5 summarizes the main findings and discusses policy implications.

2. Data

We use data for Gini coefficients from by the Standardized World Income Inequality Database (SWIID), which combines information from the United Nations World Income Database (UNWIDER) and the Luxembourg Income Study (LIS). The database provides comparable estimates of Gini indices (and associated standard deviations) of gross income inequality for 173 countries for as many years as possible from 1960 to 2010.[5] Gini coefficients are theoretically bounded between 0 (each reference unit receives an equal share of income) and 100 (a single reference unit receives all income while all the others receive nothing). In our sample they range from 18 to 78, with higher levels of inequality typically recorded for low and middle-income countries (Table 1).

The measure of financial globalization used in this paper is based on a *de jure* indicator of capital account restrictions. While *de jure* measures are noisy indicators of the true degree of openness of the capital account, they have the advantage of being less sensitive to reverse causality issues in panel regressions (Collins, 2007). Data for capital account openness are taken from the Chinn and Ito (2008) database. While alternative *de jure* measures of capital account openness have been proposed in the literature (e.g. Quinn and Toyoda, 2008), the Chinn and Ito index (Kaopen) provides the largest country and time period coverage.[6] The index measures a country's degree of capital account openness based on the binary dummy variables that codify the tabulation of restrictions on cross-border financial transactions reported in the IMF's *Annual Report on Exchange Arrangements and Exchange Restrictions (AREAER)* database.[7] The index is available for an unbalanced panel of 182 countries from 1970 to 2010, and it ranges from −1.9 (more restricted capital account) to 2.5 (less restricted). The score of the capital account openness index varies greatly across income groups, with higher restrictions typically recorded in low-income and lower-middle income countries (Table 1).

We supplement the data on the Gini coefficients with data on top

[2] The empirical evidence on the effect of financial development on inequality both based on cross-country evidence and on the US is mixed. Beck et al. (2007), using a sample of 65 countries over the period 1960–2005, report a negative relationship between financial development and the growth rate of the Gini coefficient. In contrast, Jauch and Watzka (2012), by extending the sample to 138 countries for the years 1960–2008, find that financial increases inequality when controlling for unobservable country-specific factors (country fixed effects). Similarly, de Haan and Sturm (2016) find that financial liberalization increases inequality. Beck et al. (2010) find that bank deregulation in the US reduced inequality by boosting income for people in the lower part of the distribution but has little impact on income above the median. In contrast, Jerzmanowski and Nabar (2013) show that financial market development contributed to the rise in the skill premium and residual wage inequality in the United States since the 1980s.

[3] Larrain (2015), using sectoral data for a sample of 20 advanced economies, finds that capital account openness increases wage inequality, particularly in industries with high financial dependence. Das and Mohapatra (2003), using a panel data of 11 emerging market economies, find that the capital account liberalization reforms introduced in these countries between 1986 and 1996 led to an increase in inequality by boosting income for people in the top quintile of the distribution, while having little effect on other income categories. Bumann and Lensink (2016) present a theoretical model in which they consider changing reserve requirements and opening up to foreign capital as alternate ways of liberalization of the financial sector. In their model, the impact of liberalization on inequality is ambiguous; empirically they find that the impact is positive and depends on the level of financial development, which is consistent with our findings.

[4] Another channel for the distributional effects of capital account liberalization is the following: since capital and skilled labor tend to be complements (Cragg and Epelbaum, 1996), opening the capital account to flows of foreign direct investment (FDI) can increase the demand for skilled labor compared to unskilled labor, leading to higher wage inequality. At the macro level, it is often difficult to differentiate the effect of FDI inflows from that of portfolio and debt flows given the high correlations between these flows. A careful analysis would require using sector level data, as done in Larrain (2015) for some advanced economies; such sectoral data are not available for the vast majority of emerging and low-income economies.

[5] See Solt (2009) for details on the methodology.

[6] While Kaopen is used as a baseline, alternative measures of capital account openness are also considered as a robustness check (see Section 3).

[7] See Chinn and Ito (2008) for details on the methodology.

Table 1
Descriptive statistics by income groups.

	N	Average	SD	Min	Max
Panel A. All countries					
Gini	4334	44.531	9.274	17.590	77.965
D.Gini	4020	−0.014	1.836	−13.567	19.571
Kaopen	6023	−0.002	1.529	−1.856	2.456
D.Kaopen	5829	0.024	0.370	−3.253	3.253
Panel B. High income					
Gini	1542	42.653	6.601	25.022	64.877
D.Gini	1464	0.058	1.716	−13.567	10.676
Kaopen	1667	1.036	1.516	−1.856	2.456
D.Kaopen	1618	0.044	0.299	−2.292	2.292
Panel C. Upper middle income					
Gini	1298	45.699	10.692	17.590	77.965
D.Gini	1187	−0.011	1.867	−11.059	10.844
Kaopen	1538	−0.138	1.470	−1.856	2.456
D.Kaopen	1488	0.023	0.449	−3.253	2.556
Panel D. Lower middle income					
Gini	937	44.991	9.533	23.568	77.480
D.Gini	863	−0.054	1.939	−8.646	19.571
Kaopen	1606	−0.352	1.342	−1.856	2.456
D.Kaopen	1551	0.014	0.384	−3.253	3.253
Panel D. Low income					
Gini	557	46.235	10.656	25.146	75.853
D.Gini	506	−0.161	1.912	−8.706	6.917
Kaopen	1212	−0.793	1.017	−1.856	2.456
D.Kaopen	1172	0.011	0.323	−1.935	2.988

Table 2
Number of capital account liberalization reforms.

	70s	80s	90s	2000s	1970–2010
All	38	25	100	61	224
High income	15	7	23	14	58
Upper middle income	11	9	28	31	79
Lower middle income	5	6	31	12	54
Lower income	7	3	18	5	33

incomes shares from Atkinson et al. (2011), poverty rates from the World Bank, and labor shares of income from the UN National Accounts.

The stylized facts emerge from this descriptive evidence are that both inequality and capital account openness have increased over the past two decades. Inequality has increased more persistently in high income countries. Much of the increase in capital account openness occurred during the 1990s, also the period of the largest increase in income inequality.

Examining the behavior of inequality before and after the removal of restrictions on the capital account requires information about the date on which the restrictions were lifted. This information is difficult to obtain for a large set of countries, as ideally it would require information on dates of policy decrees or legislative changes. To infer the timing of major policy changes, we identify capital account liberalization episodes by assuming that a liberalization takes place when, for a given country at a given time, the annual change in the Kaopen indicator exceeds by two standard deviations the average annual change over all observations (i.e. exceeds 0.76). [8] This criterion identifies 224 episodes of liberalization (Table 2), which are associated with an increase in the Kaopen indicator that ranges from 0.70 to 3.3, and with an average increase of about 1.3. Most liberalization episodes have occurred during the last two decades, particularly during the 1990s. Examples of these liberalization episodes include those of many advanced European countries in 1993—that is after the completion of the Single Market.

[8] A similar strategy has been followed in previous papers to identify episodes of stock market liberalizations (Henry, 2007) and labor and product market reforms (Bernal-Verdugo et al., 2013; Bouis et al., 2012).

3. The distributional effects of liberalization

This section examines the effects of capital account liberalization reforms on inequality. Before turning to the empirical evidence, it is useful to look at whether capital account liberalization episodes have been followed by an increase in inequality.

Descriptive statistics on the change in the Gini coefficient before and after the beginning of these liberalization episodes suggest that capital account liberalizations, on average, have been typically associated with an increase in the Gini coefficient of about 0.8 percentage point (2 percent) in the short term—in the year after the occurrence of a liberalization episode—and of about 1.2 percentage points (2½ percent) in the medium term—5 years after the occurrence of a liberalization episode (Fig. 1). The rest of the section checks whether this descriptive evidence holds up to more formal tests.

3.1. Methodology

To assess the impact of capital account liberalization on inequality, we follow the Autoregressive Distributed Lag (ARDL) approach of Cerra and Saxena (2008) and Romer and Romer (2010), among others. This approach is particularly suited to assess the dynamic response of the variable of interest in the aftermath of a reform (a capital account liberalization episode in our case). The methodology consists of estimating a univariate autoregressive equation and deriving the associated impulse response functions:

$$g_{it} = a_i + \gamma_t + \sum_{j=1}^{l} \beta_j g_{i,t-j} + \sum_{j=0}^{l} \delta_j D_{i,t-j} + \sum_{j=1}^{l} \vartheta_j X_{i,t-j} + \varepsilon_{it} \tag{1}$$

where g is the annual change in the (log of the) Gini coefficient; D is a dummy variable which is equal to 1 at the start of a capital account liberalization episode and zero otherwise; a_i are country fixed effects included to control for unobserved cross-country heterogeneity in inequality and also to control for the fact that in some countries inequality is measured using income data while in other countries using consumption data; γ_t are time fixed effects to control for global shocks.

We include lagged inequality growth to control for the normal dynamics of inequality. In addition, since the variables affecting inequality in the short term are typically serially correlated, this also helps to control for various factors that may influence inequality.

Finally, since several types of economic reforms are often implemented simultaneously—this is particularly the case for current account and capital account reforms—we include a set of other structural reform variables (X) to distinguish the effect of capital account liberalization episodes from others. Specifically, the set of reform variables included as controls are: (i) current account reforms, defined as an episode where the annual change of the Quinn and Toyoda (2008) measure of current account openness exceeds by two standard deviations the average annual change over all observations; and (ii) regulation reforms, defined as an episode where the annual change in a composite measure of credit, product and labor market regulation exceeds by two standard deviations the average annual change over all observations.[9]

Equation (1) is estimated using OLS on an unbalanced panel of annual observations from 1970 to 2010 for 149 advanced and developing economies. While the presence of a lagged dependent variable and country fixed effects may in principle bias the estimation of δ_j and β_j in small samples (Nickell, 1981), the length of the time dimension mitigates this concern.[10] The number of lags chosen is 2, but different lag lengths are tested as a robustness check (see next section).

[9] The data come from the Fraser Institute's Economic Freedom of the World database. Higher values of the indicators indicate more open and competitive markets.
[10] The finite sample bias is in the order of 1/T, where T in our sample is 41. Robustness checks using a two-step system-GMM estimator are provided in the next section.

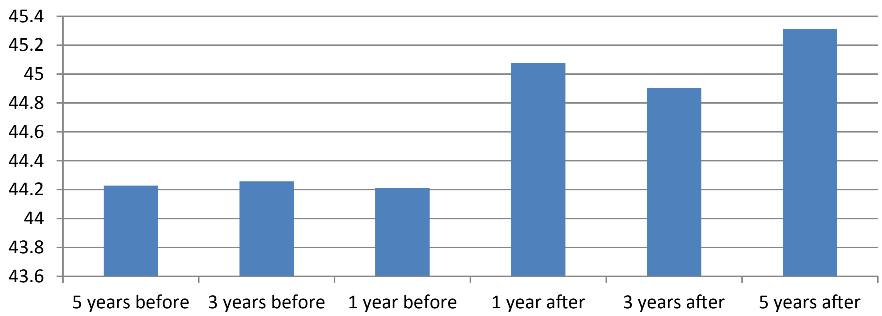


Fig. 1. The evolution of inequality before and after capital account liberalizations, absolute changes in Gini.

Impulse response functions (IRFs) are used to describe the response of inequality following a capital account liberalization episode. The shape of these response functions depends on the value of the δ and β coefficients; For instance, the simultaneous response is δ_0, the one-year ahead cumulative response is $\delta_0 + (\delta_1 + \beta_0\delta_0)$. The confidence bands associated with the estimated impulse-response functions are obtained using the estimated standard errors of the estimated coefficients, based on clustered (at country-level) heteroskedasticity robust standard errors.

Table 3
The effect of capital account liberalization on inequality (1970–2010), OLS.

	(I) Gini growth	(II) Capital flow, percent of GDP	(III) Change in top 1% income share	(IV) Change in top 5% income share	(V) Change in top 10% income share
Dep. variable (t-1)	0.272***	0.191***	−0.261**	−0.241*	−0.187*
	(4.52)	(3.88)	(-2.19)	(-1.83)	(-1.68)
Dep. variable (t-2)	0.127***	0.044	−0.036	−0.007	−0.024
	(2.77)	(1.26)	(-0.50)	(-0.12)	(-0.39)
Capital account reform (t)	0.766***	4.819**	0.193	0.847**	0.815
	(3.12)	(2.22)	(0.65)	(2.34)	(1.53)
Capital account reform (t-1)	0.143	−0.687	0.358**	0.672**	0.814*
	(0.45)	(-0.43)	(2.02)	(2.06)	(1.94)
Capital account reform (t-2)	−0.048	−1.826	0.538**	0.826*	1.334**
	(-0.16)	(-1.12)	(2.15)	(1.95)	(2.09)
Current account reform (t)	0.285	0.841	0.234	0.108	0.429
	(1.00)	(0.50)	(1.32)	(0.35)	(1.11)
Current account reform (t-1)	0.135	1.392	−0.083	0.192	0.012
	(0.43)	(0.54)	(-0.49)	(1.00)	(0.04)
Current account reform (t-2)	0.579*	0.123	−0.150	−0.094	0.080
	(1.65)	(0.01)	(-1.37)	(-0.46)	(0.31)
Regulation reform (t)	−0.333	2.372	−0.024	0.080	0.665
	(-1.05)	(1.15)	(-0.06)	(0.15)	(0.98)
Regulation reform (t-1)	−0.414	−2.837	0.918*	1.200	0.899
	(-0.96)	(-1.14)	(1.71)	(1.60)	(1.09)
Regulation reform (t-2)	−0.059	−1.274	1.482	1.850	1.930
	(-0.18)	(-0.69)	(1.17)	(1.23)	(0.96)
N	2071	2242	429	365	375
R^2	0.21	0.57	0.20	0.22	0.18

Note: T-statistics based on robust clustered standard errors in parenthesis. ***,**,* denote significance at 1 percent, 5 percent and 10 percent, respectively. Capital account reforms are identified as episodes when, for a given country at a given time, the annual change in the Kaopen indicator exceeds by two standard deviations the average annual change over all observations. Current account reforms are identified as episodes when, for a given country at a given time, the annual change in the Quinn and Toyoda current account indicator exceeds by two standard deviations the average annual change over all observations. Regulatory reforms are identified as episodes when, for a given country at a given time, the annual change in the EFW regulatory indicator exceeds by two standard deviations the average annual change over all observations.

Since some of the observations of the dependent variable are based on estimates, the regression residuals can be thought of as having two components. The first component is sampling error (the difference between the true value of the dependent variable and its estimated value). The second component is the random shock that would have been obtained even if the dependent variable was observed directly as opposed to estimated. This would lead to an increase in the standard deviation of the estimates and lower the *t*-statistics. To address this issue, and as a further robustness check, equation (1) is also been estimated with Weighted Least Squares (WLS). Specifically, the WLS estimator assumes that the errors ε_i in equation (1) are distributed as $\varepsilon_i \sim N(0, \sigma^2/s_i)$, where s_i are the estimated standard deviations of the Gini coefficient for each country i provided in the SWIID database, and σ^2 is an unknown parameter that is estimated in the second-stage regression.

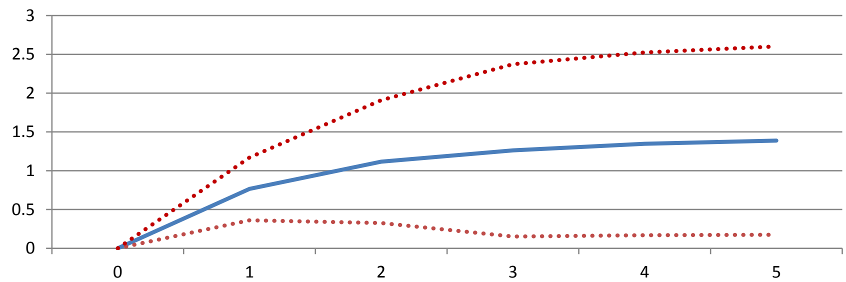
3.2. *Results*

Our baseline regression is reported in Table 3, column I. Using these estimates, we trace out the response of capital account liberalization reforms on inequality in Fig. 2. The figure presents the estimated effect of liberalization and the associated 90 percent confidence bands (dotted lines). Capital account liberalization episodes have statistically significant and long-lasting effects on income inequality: the Gini index increases by about 0.8 percent in the very short term—1 year after the occurrence of the reform episode—and by about 1.4 percent in the medium term—5 years after the occurrence of the reform episode. This medium-term effect is equivalent to an increase in the average Gini coefficient in the sample from about 44.5 to about 45.3. This effect is not negligible given that the Gini coefficient changes fairly slowly over time over time—the effect corresponds to approximately ½ standard deviation of the average change of the Gini in the sample.

The estimated coefficients reported in Table 3 indicate that capital account liberalization has a contemporaneous effect on inequality (i.e. within the year); moreover, they suggest that the persistent effect of liberalization of on inequality could be driven by the high degree of persistence of inequality itself. Both these aspects of the results deserve further scrutiny.

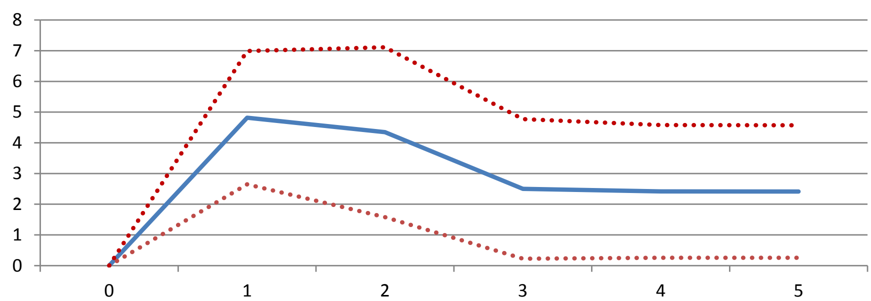
To probe whether a contemporaneous impact is plausible, we estimate an equation similar to equation (1) but with the share of capital flows—defined as the changes in total asset and liabilities—in GDP as the dependent variable.[11] The idea is that if capital account liberalization indeed has a contemporaneous effect on inequality, we should then observe a marked increase in capital flows following a reform.[12] The results of this exercise are reported in Fig. 3 (and column II, Table 3), and suggest that capital account liberalization reforms are in fact associated with a contemporaneous increase in capital flows of about 5 percent of

[11] Data on assets and liabilities are taken from Lane and Milesi-Ferretti (2007).
[12] We thank an anonymous referee for suggesting this analysis.



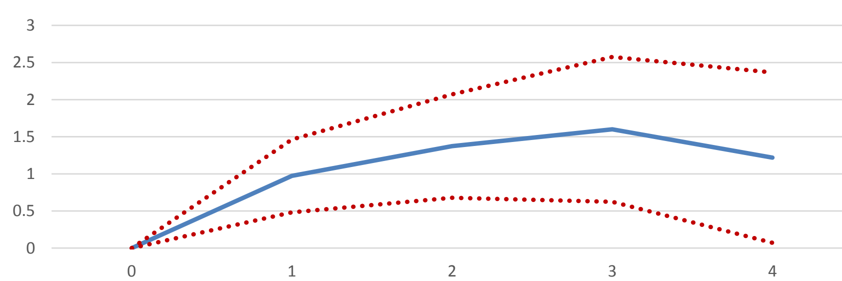
Note: IRFs are estimated using the specification in equation (1). The solid line corresponds to the IRF; dotted lines correspond to 90 percent confidence bands. The x-axis denotes time. t=0 is the year of the reform.

Fig. 2. The effect of capital account liberalization on inequality, Gini (percent).



Note: IRFs are estimated using the specification in equation (1). The solid line corresponds to the IRF; dotted lines correspond to 90 percent confidence bands. The x-axis denotes time. t=0 is the year of the reform.

Fig. 3. The effect of capital account liberalization on capital flows, percent of GDP.



Note: IRFs are estimated using the specification in equation (2). The solid line corresponds to the IRF; dotted lines correspond to 90 percent confidence bands. The x-axis denotes time. t=0 is the year of the reform.

Fig. 4. The effect of capital account liberalization on inequality, local projections, Gini (percent).

GDP. The effect is persistent but as in the case of inequality the persistent effect is mostly driven by the high degree of persistence of flows (column II, Table 3).

Next, to shed some light on the source of the persistent effect of liberalization of inequality, we use an alternate econometric specification, the local projection method proposed by Jordà (2005). With this method, the dynamic response of inequality to capital account liberalization reforms is not obtained through the estimated coefficients of lagged inequality and lagged capital account liberalization reforms—which may be highly correlated, and therefore lead to weak t-statistic for the individual coefficients—but by tracing directly the evolution of inequality in the aftermath of a capital account liberalization episode. In particular, the following equation is estimated for each $k = 1, ..4$:

$$lnGini_{i,t+k} - lnGini_{i,t} = a_i^k + \gamma_t^k + \sum_{j=1}^{l} \beta_j g_{i,t-j} + \delta^k D_{i,t} + \vartheta^k X_{i,t} + \varepsilon_{i,t}^k \quad (2)$$

where δ^k is the estimated response of inequality in each period $t + k$ to a

Table 4
The effect of capital account liberalization on inequality (1970–2010), Local projection methods.

	K = 1	K = 2	K = 3	K = 4
Capital account reform (t)	0.972***	1.374***	1.600***	1.219*
	(3.26)	(3.25)	(2.70)	(1.75)
Current account reform (t)	0.184	0.355	1.077	2.599***
	(0.65)	(0.74)	(1.48)	(2.74)
Regulation reform (t)	−0.296	−0.803	−1.070	0.157
	(-0.75)	(-1.02)	(-0.90)	(0.10)
N	2071	2071	2071	2071
R^2	0.23	0.26	0.23	0.27

Note: Estimates based on equation (2). K = 1, ..,4 denotes the year following the capital account liberalization reform. T-statistics based on robust clustered standard errors in parenthesis. ***,**,* denote significance at 1 percent, 5 percent and 10 percent, respectively. Capital account reforms are identified as episodes when, for a given country at a given time, the annual change in the Kaopen indicator exceeds by two standard deviations the average annual change over all observations. Current account reforms are identified as episodes when, for a given country at a given time, the annual change in the Quinn and Toyoda current account indicator exceeds by two standard deviations the average annual change over all observations. Regulatory reforms are identified as episodes when, for a given country at a given time, the annual change in the EFW regulatory indicator exceeds by two standard deviations the average annual change over all observations.

capital account liberalization reform introduced at time *t*, and all the other terms have the same interpretation as in equation (1).[13] The results obtained by estimating equation (2) are similar to those from the ARDL approach and not statistically significantly (Fig. 4 and Table 4). They confirm that capital account liberalization reforms have statistically significant short-term and medium-term effect on inequality. In particular, capital account liberalization reforms has typically increased the Gini index by about 1.0 percent in the short term—1 year after the occurrence of the episode—and by about 1.2 percent in the medium term—5 years after the occurrence of the episode.

In sum, these two exercises provide supportive evidence that (i) capital account liberalization is associated with contemporaneous increases in capital flows, making the contemporaneous impact on inequality more plausible; (ii) liberalization has a persistent effect on inequality, and this effect is not due solely to the high persistence of the Gini coefficient.

3.2.1. Depth and direction of capital account reforms

Does the impact on inequality vary with the depth of the liberalization? To answer this question, we repeat the empirical analysis by identifying episodes based on different thresholds, specifically, one and three standard deviations of the average annual change over all observations (instead of the two standard deviations threshold used in the baseline results). The results remain statistically significant for these alternative thresholds and the magnitude of the effect on inequality does indeed increase with the depth of capital account liberalization (Fig. 5).

Another interesting question is whether capital account *restrictions* reduce inequality. To answer this, we construct episodes when, for a given country at a given time, the annual change in the Kaopen indicator is two standard deviations *below* the average annual change over all observations.[14] The results of this exercise show that while capital account restrictions tend to reduce inequality, the effect is not statistically significantly different from zero (Fig. 6).

3.3. Robustness checks

3.3.1. Measurement errors

The significance of our results could be affected by the quality of the data and the fact that some observations of the dependent variables have themselves been estimated. To gauge the extent of this problem, we re-estimate equation (1) with WLS using as analytical weights the inverse of the standard errors associated with each year-country observation of the Gini.[15] The results are reported in Fig. 7 (Table 5, column I) and confirm that our main finding. While the WLS estimates produce similar results to those obtained with OLS, the medium-term effect is somewhat larger (about 1½ percent), but statistically not different from the baseline results.

3.3.2. Lag parameterizations

Teulings and Zubanov (2014) and Bernal-Verdugo et al. (2013) note that the IRFs using ARDL models can be sensitive to the choice of the number of lags. As a check on our results, we re-estimate equation (1) using two different lag-parameterizations, ARDL (1, 1) and ARDL (5, 5). The results reported in Fig. 8 (Table 5, columns II-III) show that the IRFs tend to be close to each other, and the differences in the IRFs are not statistically significant.

3.3.3. Different measures of capital account openness

We also test if the impact of financial globalization on inequality is robust to the use of alternative measures of capital account openness. In particular, we re-estimate equation (1) using the Quinn and Toyoda (2008) indicator of capital account openness and defining episodes of liberalization in a way similar to that in the baseline. The results obtained with this measure again point to a statistically significant and persistent impact of capital account liberalization reforms on inequality (Fig. 9, Table 5 column IV). While the short-term effect is very similar to the one reported in the baseline, the medium-term effect appears to be significantly higher (about 2½ percent—that is, about ¾ standard deviation of the average change of the Gini coefficient in the sample) than the one obtained using the Kaopen index.

3.3.4. Different measures of inequality

An important robustness check is to see whether the results are robust to different measures of inequality. Equation (1) is re-estimated using the top shares of incomes, the top 1 percent, top 5 percent, and top 10 percent.[16] The results confirm that capital account liberalization reforms tend to increase inequality using these measures (Fig. 10 and Table 3, columns III-V). In particular, liberalization has typically led to a medium-term increase in the top 1 percent income share of about 1 percent, and to a medium-term increase in the top 5 and top 10 percent income shares of about 2 percent.

3.3.5. Addressing omitted variable bias

Potential reverse causality is likely to not be an issue since the decision on whether to liberalize the capital account is unlikely to be influenced by inequality.[17] However, it could still be the case that unobserved factors influencing the dynamics of inequality over time could affect the probability of financial liberalization. This could be the case, for example, if governments that choose to liberalize capital account are more right-wing and less likely to implement redistributive policies.[18] While including reforms in other macroeconomic areas should mitigate this problem, as an addition check, we have also re-

[13] The main shortcoming of this approach compared to ARDL is that the medium- and long-term effects tend to be less precisely estimated.

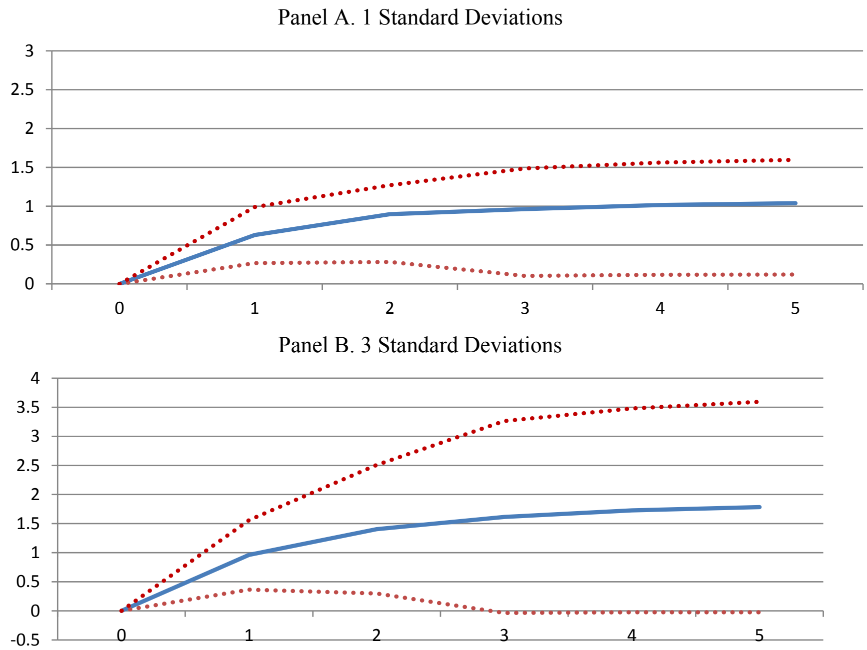
[14] According to this criterion 157 episodes of capital account restrictions are identified.

[15] The size of the standard error largely depends on data availability in the UNWIDER and LIS database. Solt (2009) reports that about 30 percent of the observations have associated standard errors of 1 point or less on the 0 to 100 scale of the Gini index. Over 60 percent of the standard errors are less than 2 points, and more than 85 percent are less than 3 points. Fewer than 3 percent of observations have standard errors greater than 5 points, and 0.3 percent of observations are greater than 10 points.

[16] Data on top income shares are taken from Atkinson et al. (2011).

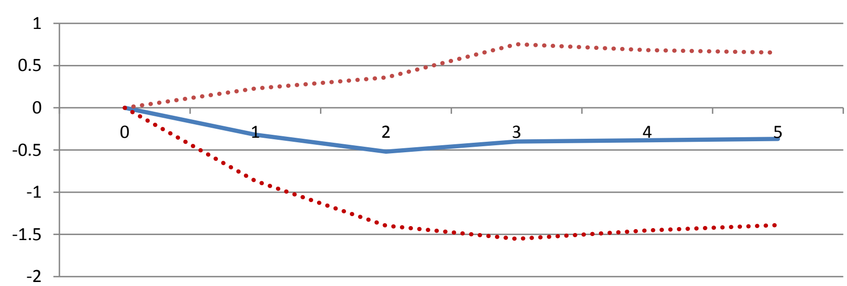
[17] Indeed, Granger causality tests suggest that lagged inequality does not significantly affect the probability of a capital account liberalization episode.

[18] We thank an anonymous referee for pointing out to this potential source of endogeneity.



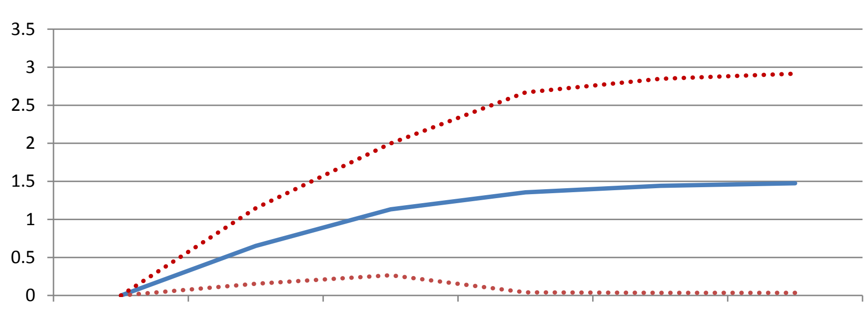
Note: IRFs are estimated using the specification in equation (1). The solid line corresponds to the IRF; dotted lines correspond to 90 percent confidence bands. The x-axis denotes time. t=0 is the year of the reform.

Fig. 5. The effect of capital account liberalization on inequality, Depth of liberalization, Gini (percent).



Note: IRFs are estimated using the specification in equation (1). The solid line corresponds to the IRF; dotted lines correspond to 90 percent confidence bands. The x-axis denotes time. t=0 is the year of the reform.

Fig. 6. The effect of capital account restriction on inequality, Gini (percent).



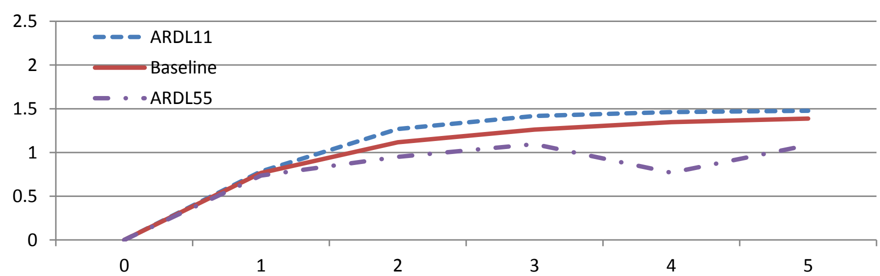
Note: IRFs are estimated using the specification in equation (1). The solid line corresponds to the IRF; dotted lines correspond to 90 percent confidence bands. The x-axis denotes time. t=0 is the year of the reform.

Fig. 7. The effect of capital account liberalization on inequality, WLS, Gini (percent).

Table 5
The effect of capital account liberalization on inequality, Robustness checks.

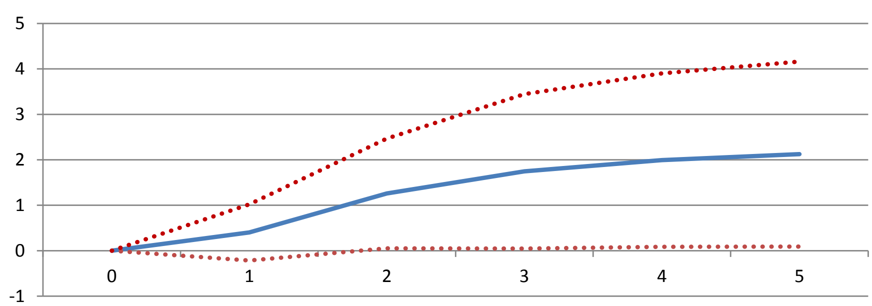
	(I)	(II)	(III)	(IV)	(v)	(VI)	(VII)
	WLS	ARDL (1,1)	ARDL (5,5)	Quinn-Toyoda	Additional controls	IV	GMM
Dep. variable (t-1)	0.331***	0.307***	0.287***	0.255***	0.255***	0.258***	0.048
	(3.74)	(4.82)	(4.75)	(4.34)	(4.70)	(5.43)	(0.62)
Dep. variable (t-2)	0.026		0.176***	0.143***	0.164***	0.136***	−0.080*
	(0.30)		(3.66)	(2.94)	(3.40)	(3.42)	(-1.70)
Capital account reform (t)	0.652**	0.783***	0.626***	0.402	0.815**	0.748**	0.924***
	(2.16)	(3.27)	(2.72)	(1.07)	(2.91)	(2.43)	(3.02)
Capital account reform (t-1)	0.265	0.244	−0.120	0.755*	0.239	0.094	0.568*
	(1.17)	(0.81)	(-0.37)	(1.65)	(0.82)	(1.44)	(1.94)
Capital account reform (t-2)	0.046		−0.070	0.211	0.173	0.030	0.218
	(0.14)		(-0.24)	(0.53)	(0.62)	(0.45)	(0.72)
N	2071	2111	1986	2137	1774	2071	2071
R^2	0.22	0.20	0.25	0.21	0.31	0.18	

Note: T-statistics based on robust clustered standard errors in parenthesis. ***,**,* denote significance at 1 percent, 5 percent and 10 percent, respectively. Capital account reforms are identified as episodes when, for a given country at a given time, the annual change in the Kaopen indicator exceeds by two standard deviations the average annual change over all observations. Controls included but not reported.



Note: IRFs are estimated using the specification in equation (1). The solid line corresponds to the IRF; dotted lines correspond to 90 percent confidence bands. The x-axis denotes time. t=0 is the year of the reform.

Fig. 8. The effect of capital account liberalization on inequality, different lags, Gini (percent).



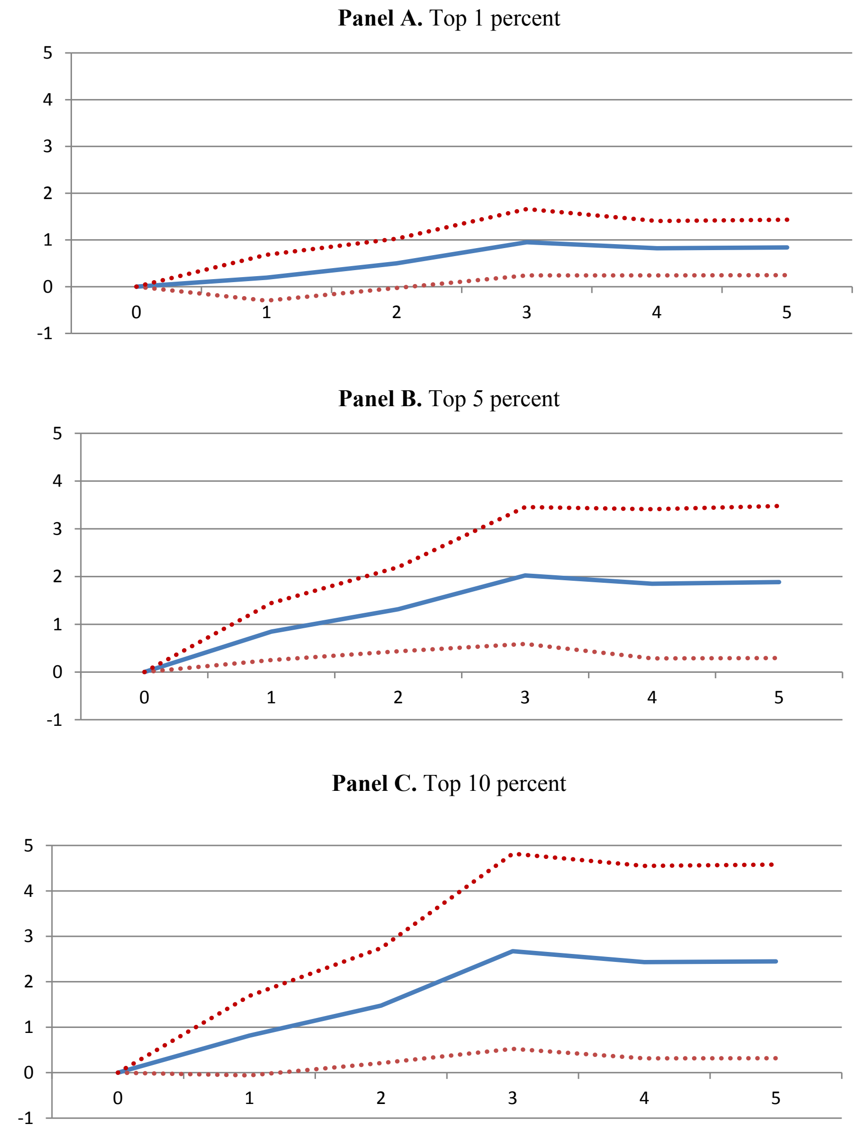
Note: IRFs are estimated using the specification in equation (1). The solid line corresponds to the IRF; dotted lines correspond to 90 percent confidence bands. The x-axis denotes time. t=0 is the year of the reform.

Fig. 9. The effect of capital account liberalization on inequality, Quinn and Toyoda measure of capital account openness, Gini (percent).

estimated equation (1) using (i) a discrete variable for left-, center-, right-wing government[19]; and (ii) changes in the share of redistributive policies—proxied by changes in the difference between gross and net Gini coefficients.

In addition to these two variables, we including a set of control variables which may affect the evolution of inequality and influence the impact of capital account liberalization, namely: (i) GDP growth; (ii) the level and the square of log GDP per capita; (iii) changes in trade openness (defined as the sum of exports and imports over GDP); (iv) changes in the

[19] The variable is taken from the Database of Political Institutions and assumes value 0 for left-wing governments, 1 for center-wing governments, and 2 for right-wing governments.



Note: IRFs are estimated using the specification in equation (1). The solid line corresponds to the IRF; dotted lines correspond to 90 percent confidence bands. The x-axis denotes time. t=0 is the year of the reform.

Fig. 10. The effect of capital account liberalization on the top income shares, percent.

share of government expenditures in GDP; (v) changes in the share of industry and agriculture value added; (vi) changes in dependency ratios; and (vii) changes in product, labor and credit market regulations.

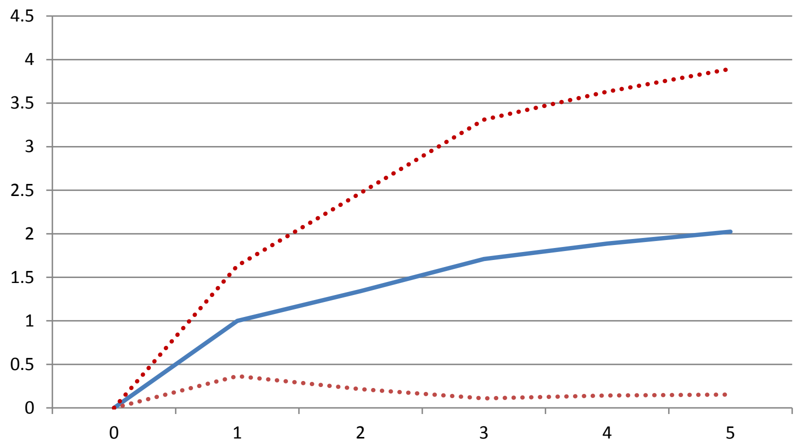
The results of this exercise are presented in Fig. 11 (Table 5, column V), and confirm a significant and persistent effect of capital account liberalization reforms on inequality. The results also suggest a larger medium-term effect than the one reported in the baseline, even though the difference is not statistically significant.[20]

3.3.6. *Addressing endogeneity bias*

To address endogeneity concerns, we have also conducted an instrumental variables (IV) approach using two instruments that capture the scope for reforms and the peer pressure to reform. The scope of capital account liberalization reform is captured by the initial stance of capital account regulation—proxied by the four-year lagged value of the capital account openness indicator. The idea is that the lower is the indicator of capital account openness, the more scope there is to reform.[21] Peer pressure is proxied by a weighted-average of current and lagged capital account liberalization episodes in other countries, where the weights are determined by strength of trade linkages between other countries and the country undertaking a capital account

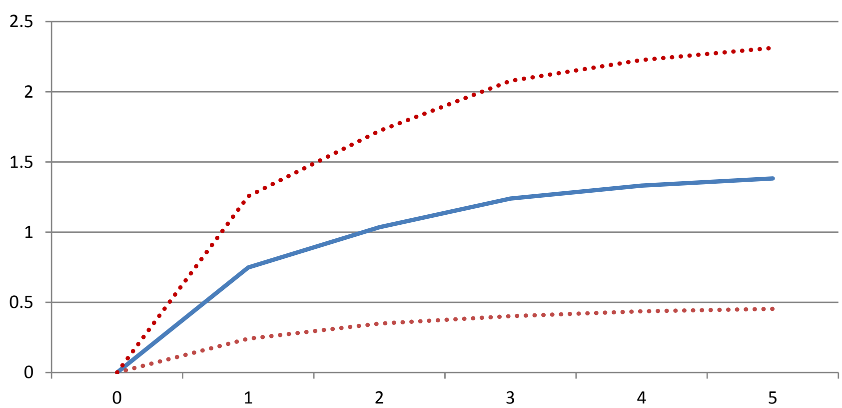
[20] Among the control variables included in the regression we find that GDP growth, the level of GDP per capita, and changes in the share of redistributive policies are positively associated with change in the Gini, while the change in the share of agriculture and the square of GDP per capita are negatively related.

[21] We use the four-year lagged level of capital account openness to mitigate the possibility that past values of the level of capital account openness may directly affect inequality. Estimates, not reported but available upon request, suggest that the effect of the *level* of capital account openness on inequality turns to be statistically insignificant after two years, once capital account liberalization reforms are considered.



Note: IRFs are estimated using the specification in equation (1). The solid line corresponds to the IRF; dotted lines correspond to 90 percent confidence bands.

Fig. 11. The effect of capital account liberalization on inequality, additional controls, Gini (percent).



Note: Note: IRFs are estimated using the specification in equation (1). The solid line corresponds to the IRF; dotted lines correspond to 90 percent confidence bands. The x-axis denotes time. t=0 is the year of the reform. The instruments are the level of the indicator of capital account liberalization lagged by four periods and current and lagged account liberalization reforms in major trading partners.

Fig. 12. The effect of capital account liberalization on inequality, instrumental variables, Gini (percent).

liberalization reform.[22] The idea is that a country is more likely to implement capital account liberalization when its main trading partners are undertaking or have undertaken capital account liberalization.

First stage estimates of capital account liberalization reforms on these instruments suggest that these are statistically significant and exhibit the expected sign.[23] In addition, both instruments can plausibly be considered as exogenous, and should not have any direct effect on the left-hand side variable. For example, reforms in other countries are not driven by outcomes in the country considered, and should not have any effect on the latter other than through pressure on domestic authorities to undertake reform.[24]

[22] We use bilateral trade weights since limited data availability precludes the construction of bilateral capital flow weights for most of the observations in the sample. For the country-time observations for which bilateral capital flows are available the correlation between bilateral trade and capital flows linkages is high (about 0.7) and statistically significant. Specifically, the instrument is computed as follows:

$$I_{i,t} = \sum_{j=1,n(j\neq 1)} D_{j,t} w_{i,j,t}$$

where $I_{i,t}$ is the instrument of capital account liberalization reform for country i, at time t ($D_{i,t}$). $D_{j,t}$ is capital account liberalization reform for country j, at time t; $w_{i,t,t}$ is the share of total export and import between country i and country j in total exports and imports for country i: $\frac{Export_{i,j,t}+import_{i,j,t}}{Export_{i,t}+import_{i,t}}$.

[23] In particular, the estimation results are the following:

$$D_{i,j} = \underset{(5.62)}{0.239} I_{i,t} + \underset{(2.77)}{0.105} I_{i,t-1} - \underset{(-6.28)}{0.010}\ Kaopen_{i,t-4}$$

with t-statics in parenthesis.

[24] The Kleibergen-Paap rk Wald F statistic of weak exogeneity (24.74) and the Hansen J statistics p-value for over-identification (0.77) suggest that these variables can be considered as strongly exogenous. In addition, estimates of the effects of these instruments on inequality are not statistically significant once episodes of capital account liberalizations are controlled for, suggesting that they do not directly affect inequality.

The results reported in Fig. 12 (Table 5, column VI) confirm a significant and persistent effect of capital account liberalization reforms on inequality. They also suggest a somewhat smaller medium-term effect than the one reported in the baseline, even though the difference is not statistically significant.

To address the possible bias in small sample due to the presence of a lagged dependent variable and country fixed effects (Nickell, 1981), equation (1) has been re-estimated using the two-step GMM estimator.[25] The results also in this case are similar and not statistically significantly different from those reported in the baseline (Fig. 13, Table 5 column VII).

3.3.7. Effect across income groups

The descriptive evidence presented in Section 2 has shown that while capital account openness has increased in all income groups, the pattern of inequality has been much more mixed. This is particularly the case during the last decade where inequality has stabilized or decreased in middle and low-income countries, while it has increased in high-income countries. This different pattern may reflect a different effect of capital account liberalization reforms on inequality across different income groups. To test for this hypothesis, we extend equation (1) to allow for a different effect across income groups. In particular, we estimate the following specification:

$$g_{it} = a_i + \gamma_t + \sum_{j=1}^{l} \beta_j g_{i,t-j} + \sum_{j=1}^{l} \vartheta_j X_{i,t-j} + \sum_{j=0}^{l} \delta_j^H D_{i,t-j} H_i + \sum_{j=0}^{l} \delta_j^M D_{i,t-j} M_i + \sum_{j=0}^{l} \delta_j^L D_{i,t-j} L_i + \varepsilon_{it} \tag{3}$$

where *H, M, L* denotes dummy for high, middle and low income countries, respectively. The results of this exercise reported in Fig. 14 (Table 6) show different effects across income groups, with the magnitude of the effect being the largest in middle-income countries, and the smallest in low income countries. While the effect in low income countries may appear small, it is not negligible given that, on average, income inequality has *declined* in these countries by more than 10 Gini percentage points over the entire sample. At the same time, while the effect is more precisely estimated for high-income countries, the effects across different income groups are not statistically different from those for the whole sample.

4. Liberalization and inequality: channels

This section tries to identify empirically some of the mechanisms through which capital account liberalization may affect inequality, namely: (i) the extent of financial development and inclusion; (ii) the occurrence of financial crises; and (iii) the impact on labor's bargaining power, which could be reflected in the labor share of income.

4.1. Financial development and inclusion

It is commonly argued that the benefits of financial globalization depend on the level of financial institutions. Kose et al. (2011) identify certain threshold levels of financial development (in particular the depth of the credit market) that an economy needs to attain before it can benefit from, and reduce the risks associated with, financial globalization. Capital account liberalization may allow better consumption smoothing and lower volatility for countries with strong financial institutions, but where institutions are weak and the access to credit is not inclusive, it may further exacerbate inequality by increasing the bias in financial access in favor of people who are well off.[26]

To test this hypothesis, we modify equation (1) by allowing the effect of capital account liberalization to vary across different degrees of financial institutions. Specifically, we estimate the following equation:

$$g_{it} = a_i + \gamma_t + \sum_{j=1}^{l} \beta_j g_{i,t-j} + \sum_{j=1}^{l} \vartheta_j X_{i,t-j} + \sum_{j=0}^{l} \delta_j^- D_{i,t-j} G(z_{it}) + \sum_{j=0}^{l} \delta_j^+ D_{i,t-j} (1 - G(z_{it})) + \varepsilon_{it} \tag{4}$$

with

$$G(z_{it}) = \frac{\exp(-\gamma z_{it})}{1 + \exp(-\gamma z_{it})}, \quad \gamma > 0,$$

in which z is an indicator of financial development, normalized to have zero mean and unit variance, and G (z_{it}) is the corresponding smooth transition function of the degree of financial development.[27] This approach is equivalent to the smooth transition autoregressive (STAR) model developed by Granger and Teravistra (1993) to assess non-linear effects above/below a given threshold or regime.[28] The main advantage of this approach relative to estimating SVARs for each regime is that it uses a larger number of observations to compute the impulse response functions of only the dependent variables of interest, improving the stability and precision of the estimates. This estimation strategy can also more easily handle the potential correlation of the standard errors within countries, by clustering at the country level.[29]

Three indicators of financial development are considered in the analysis. The first is a composite indicator of credit market freedom provided by the Fraser Institute's Economic Freedom of the World (EFW) which rates countries between 0 and 10, with higher scores being assigned to economies with deeper and more open credit markets.[30] The second indicator is the ratio of credit to GDP (Global Financial Development Database), which represents a proxy for credit market depth. The third indicator is a measure of financial inclusion and access to credit, identified as the ratio of adults in the population who have borrowed from a formal financial institution in past years (Demirguc-Kunt et al., 2015).[31]

[25] The two-step GMM-system estimates (with Windmeijer standard errors) are computed using the xtabond2 Stata command developed by Roodman. All variables are considered as endogenous (instrumented using up to 3 lags, and the two set of instruments used in the IV approach). Consistency of the two-step GMM estimates has been checked by using the Hansen and the Arellano-Bond tests. The Hansen J-test of over-identifying restrictions, which tests the overall validity of the instruments by analyzing the sample analog of the moment conditions used in the estimation process, cannot reject the null hypothesis that the full set of orthogonality conditions are valid (the p-value is 0.641).

[26] We find that the measures of the quality of financial institutions used in the paper are negatively related to the probability of financial crises.

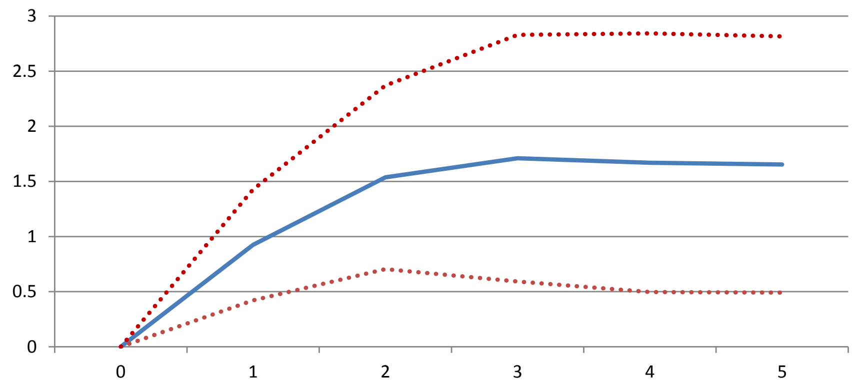
[27] γ is chosen equal to 1.5 (see Abiad et al., 2015), but the results are robust to different parameterizations.

[28] The approach is similar to estimating the effects of capital account openness on inequality based on given thresholds of financial development (such as the average or the median in the sample). An alternative approach to estimate non-linear effects would be to include an interaction term between capital account liberalization reforms and the level of financial development. That alternate approach yields similar results to the ones shown here. For example, the F-test of non-linear short-term effects based on credit-to-GDP is 4.98 (significant at 2 percent); and 4.15 (significant at 4 percent) based on financial inclusion. We found the method used here simpler for presentational purposes.

[29] This approach has been applied to model non-linearities in number of different economic issues such as exchange rates dynamics (Sarno and Taylor, 2002); sectoral performance during the business cycle (Fok et al., 2005); money demand (Chen and Wu, 2005) fiscal multipliers (Auerbach and Gorodnichenko, 2013).

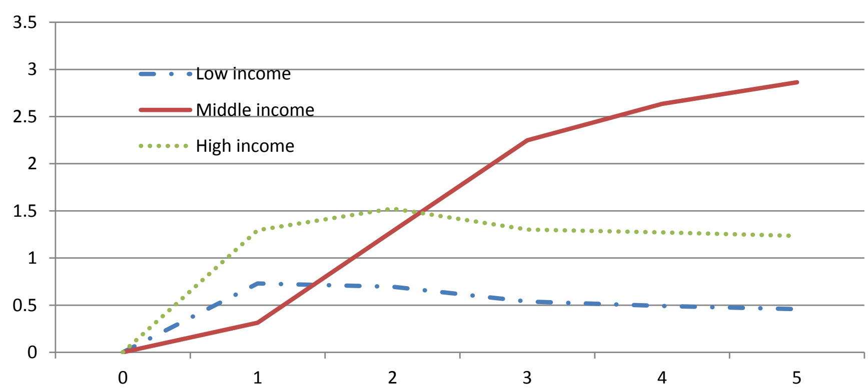
[30] The indicator is based on the following sub-components: i) Ownership of banks; ii) Foreign bank competition; iii) Private sector credit; and iv) Interest rate controls. The indicator is available for an unbalanced panel of 122 countries from 1980 to 2010, at 5-year frequency from 1980 to 2000 and at annual frequency afterward. Missing data during the five years in which annual observations are not available have been interpolated using a linear trend. The rationale for using this indicator, instead of others such as those provided by the World Bank Governance Indicator, is to maximize the country/time sample coverage.

[31] Since this indicator is only available for few years, the interaction terms have been constructed by multiplying the reform dummies by the average level of the indicator in each country.



Note: Note: IRFs are estimated using the specification in equation (1). The solid line corresponds to the IRF; dotted lines correspond to 90 percent confidence bands. The x-axis denotes time. t=0 is the year of the reform. The two-step GMM-system estimates (with Windmeijer standard errors) are computed using the xtabond2 Stata command developed by Roodman (2009a). All variables are considered as endogenous (instrumented using up to 3 lags, and the two set of instruments used in the IV approach).

Fig. 13. The effect of capital account liberalization on inequality, two-step GMM, Gini (percent).



Note: IRFs are estimated using the specification in equation (3). The x-axis denotes time. t=0 is the year of the reform.

Fig. 14. The effect of capital account liberalization on inequality across income groups, Gini (percent).

Starting with the EFW's composite indicator, the results obtained by estimating equation (3) show the effect of capital account openness on inequality depends on the level of credit market institutions, with the medium-term effect being (statistically significantly) smaller in countries with a high level of credit market openness. This result is illustrated in Panel A of Fig. 15, which presents the baseline results together with the IRFs obtained estimating equation (4) for the two degree of regimes.

The analysis is then repeated using the share of private credit to GDP. The results presented in Panel B of Fig. 15 show that the effect of capital account reforms on inequality also decreases with the depth of the credit market, with the medium-term effect of capital account liberalization reform being (statistically significantly) smaller in countries with a high level of credit market openness. Interestingly, the results suggest that in countries with very high credit-to-GDP ratio the medium-term effect of capital account liberalization on inequality is negative, even though not statistically different from zero.

In addition, we find that financial inclusion plays a significant role in shaping the response of inequality to capital account reforms, particularly over the medium term (Panel C of Fig. 15). Specifically, the figure shows that while liberalization reforms in countries with relatively low levels of financial inclusion are associated with a medium-term increase in inequality of more than 3 percent—that is, about 1 standard deviation of the average change of the Gini coefficient in the sample, in countries with relatively high levels of financial inclusion inequality increases by less than 0.1 percent over the medium term.

Financial inclusion also plays a role in determining the impact of capital account liberalization on poverty rates (Fig. 16). Though liberalization lowers the poverty ratio on average, this effect is negated in cases where financial inclusion is low.[32]

4.2. Crises

As noted in the introduction, a channel through which capital account liberalization reforms may increase income inequality is by increasing the likelihood of financial crises. To test for this hypothesis, we construct a dummy variable for those capital account liberalization episodes that have been followed by the occurrence of a financial crisis over a time horizon of 5 years—the same time horizon of the IRFs presented in earlier

[32] The literature on the relationship between capital account liberalization and the poverty rate is scant. Arestis and Caner (2010) found no statistically significant relationship between the two.

Table 6
The effect of capital account liberalization on inequality across income groups.

Dep. variable (t-1)	0.714***
	(4.55)
Dep. variable (t-2)	0.128***
	(2.77)
Capital account reform (t)*High income	1.295***
	(3.18)
Capital account reform (t-1)* High income	−0.122
	(-0.21)
Capital account reform (t-2)* High income	−0.450
	(-0.94)
Capital account reform (t)*Middle income	0.314
	(0.67)
Capital account reform (t-1)* Middle income	0.883**
	(2.19)
Capital account reform (t-2)* Middle income	0.664*
	(1.79)
Capital account reform (t)*Low income	0.730**
	(1.89)
Capital account reform (t-1)* Low income	−0.231
	(-0.44)
Capital account reform (t-2)* Low income	−0.242
	(-0.45)
N	2071
R^2	0.22

Note: T-statistics based on robust clustered standard errors in parenthesis. ***,**,* denote significance at 1 percent, 5 percent and 10 percent, respectively. Capital account reforms are identified as episodes when, for a given country at a given time, the annual change in the Kaopen indicator exceeds by two standard deviations the average annual change over all observations. Controls included but not reported. Estimated based on equation (3).

figures. The financial crisis can be either a banking, currency, or debt crisis, and the dates are identified based on the information provided by Laeven and Valencia (2010). Equation (1) is then augmented by this dummy variable, C_{it}:

$$g_{it} = a_i + \gamma_t + \sum_{j=1}^{l} \beta_j g_{i,t-j} + \sum_{j=1}^{l} \vartheta_j X_{i,t-j} + \sum_{j=0}^{l} \delta_j^{crisis} D_{i,t-j} C_{it} + \sum_{j=0}^{l} \delta_j^{no-crisis} D_{i,t-j}(1 - C_{it}) + \varepsilon_{it} \quad (5)$$

The results of this exercise show that the effect of financial globalization on inequality varies markedly between crisis and non-crisis reform episodes (Fig. 17). In particular, while crisis reform episodes are associated with a medium-term increase in inequality of more than 3.5 percent—that is, slightly more than 1 standard deviation of the average change of the Gini coefficient in the sample, in the aftermath of non-crisis reform episodes inequality increases by about 1 percent over the medium term. The difference in the IRFs increases over time, and it becomes statistically significant after the third year following a reform episode.[33]

4.3. Bargaining power and labor's share of income

To the extent that capital liberalization represents a credible threat to reallocate production abroad, it may lead to an increase in the profit-wage ratio and to a decrease in the labor share of income (Jayadev, 2007). Hence, another way to look at the distributional consequences of capital account liberalization is to examine the impact on the functional distributional of income between capital and labor. Looking at factor shares involves comparing returns to the activity of labor (the main source of income for the vast majority of the population) versus the returns to ownership (a more important source of income for the wealthy). This classification provides another perspective of how the benefits of financial globalization are shared; it also addresses the bias in measures of inequality such as the Gini which typically omits sources of income for the very wealthy.

To test if this channel is operative, we have estimated a modified version of equation (1):

$$\Delta L_{it} = a_i + \gamma_t + \sum_{j=1}^{l} \beta_j \Delta L_{i,t-j} + \sum_{j=0}^{l} \delta_j D_{i,t-j} + \sum_{j=1}^{l} \vartheta_j X_{i,t-j} + \varepsilon_{it} \quad (6)$$

where L is the labor share of income computed as the ratio of compensation of employees to GDP.[34]

The results obtained from estimating equation (6) are presented in Fig. 18. Looking at the figure it can be noted that capital liberalization episodes have statistically significant and long-lasting effects on the labor share of income. In particular, the estimates suggest that reforms have typically decreased the labor share of income by about 0.6 percentage point in the short term—1 year after the reform—and by about 0.8 percentage point in the medium term—5 years after the reform. This result is consistent with Jayadev (2007), who reports an effect of capital account openness on the labor share of income ranging between 0.5 and 1 percentage point.

Similarly, repeating the analysis for the labor share of income we find that capital account liberalization reforms tend to have the largest medium-term effects on high and middle income countries, while the effect on the low-income group countries is not statistically significant. This result is consistent with previous empirical evidence suggesting that the impact of international financial flows on inequality and labor market shares tends to be larger in advanced economies (Jaumotte et al., 2013; Jayadev, 2007).

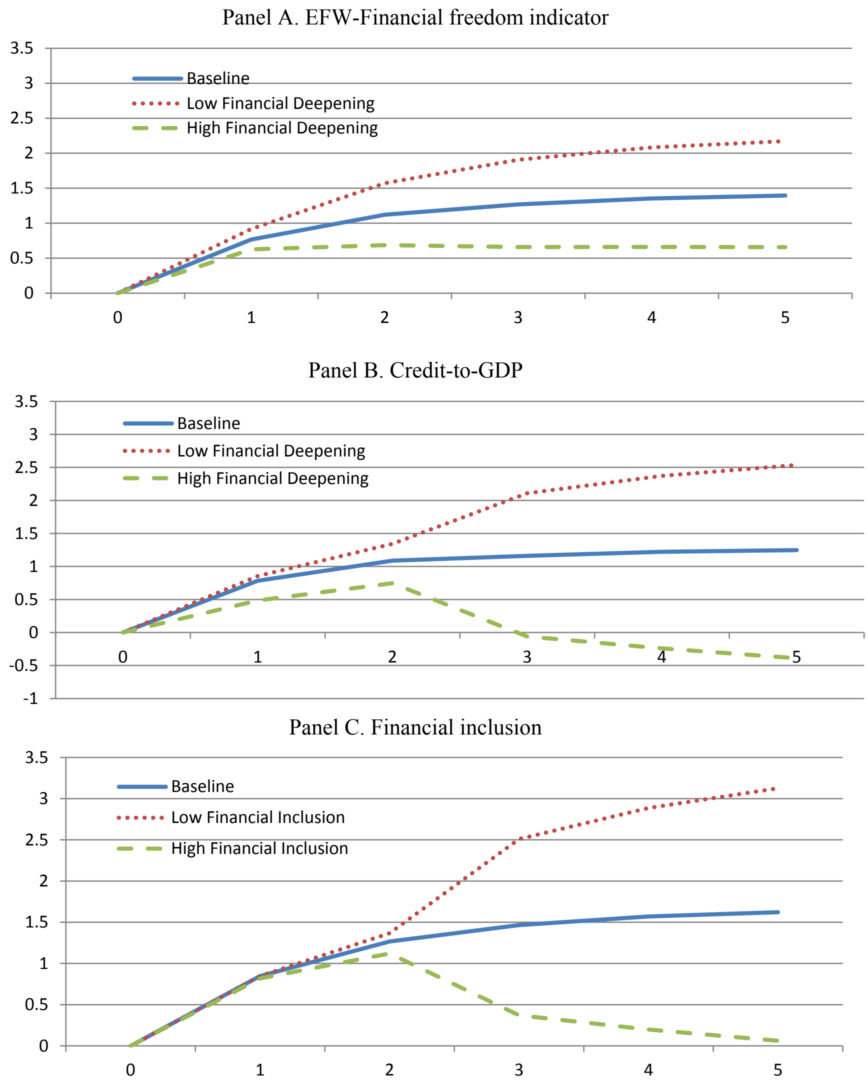
5. Conclusions

In theory, financial globalization can generate an array of benefits that boost long-run growth and welfare. However, whether these possible benefits are typically shared across all segments of the population has not been a subject of much study. The aim of this paper is to fill this gap through a comprehensive study of the distributional impacts of capital account liberalization. Using a panel of 149 countries and data covering 1970 to 2010, we find that capital account liberalization episodes are associated with a statistically significant and persistent increases in the Gini measure of inequality and in top income shares. In particular, we find that, on average, capital account liberalization reforms have typically increased the Gini coefficient by about 0.8 percent in the short term (1 year after the occurrence of the liberalization reform) and by about 0.7–3½ percent in the medium term (5 years after). These effects are economically significant given that the Gini coefficient changes very slowly over time—they correspond to approximately ½-1 standard deviation of the average change of the Gini in the sample.

This finding does not imply that countries should not undertake capital account liberalization, but it suggests an additional reason for caution. Countries where reduction in inequality is an important policy goal may need to design and sequence liberalization in a manner that balances this consideration against the other effects. Our results provide some guidance on the appropriate design and sequencing. We find that the occurrence of crises and the level of financial development and

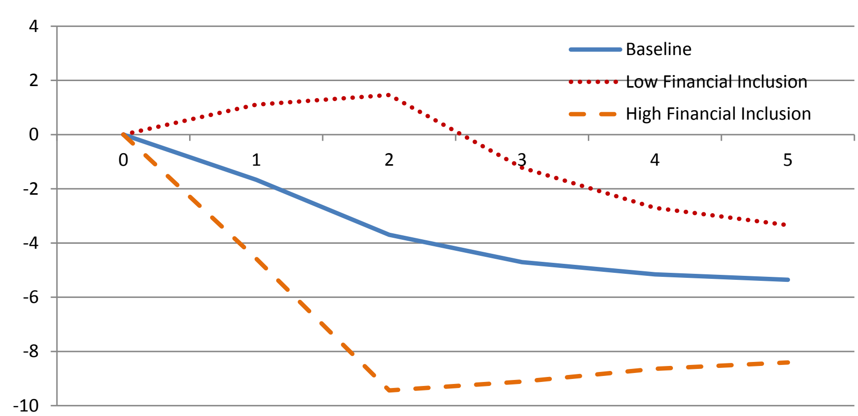
[33] In results not reported here, we find that financial crises *per-se* are associated with a significant and long-lasting increase in inequality. In particular, the estimates suggest that financial crises have typically increased the Gini index by about 0.1 percent in the short term—1 year after the occurrence of the reform episode—and by about 2.5 in the medium term—5 years after the occurrence of the crisis. Including financial crises as a separate variable does not affect our main results.

[34] Data are taken from the detailed aggregate tables of the UN national accounts, table 203 using the SNA 1993 methodology. Where multiple series were available (since the UN collects data using multiple methods), we apply the first difference of the labor share from the later series to the labor share derived from the earlier series. One shortcoming of our measure is that it does not include the labor income part of the income of self-employed.



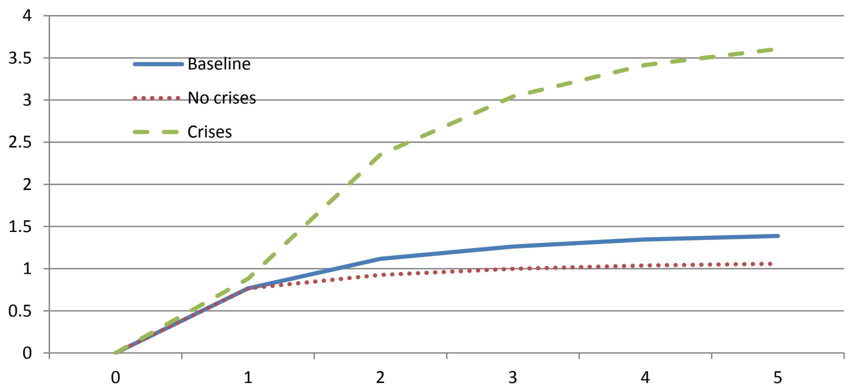
Note: IRFs are estimated using the specification in equation (4). The x-axis denotes time. t=0 is the year of the reform.

Fig. 15. The effect of capital account liberalization on inequality, the role of financial institutions, Gini (percent).



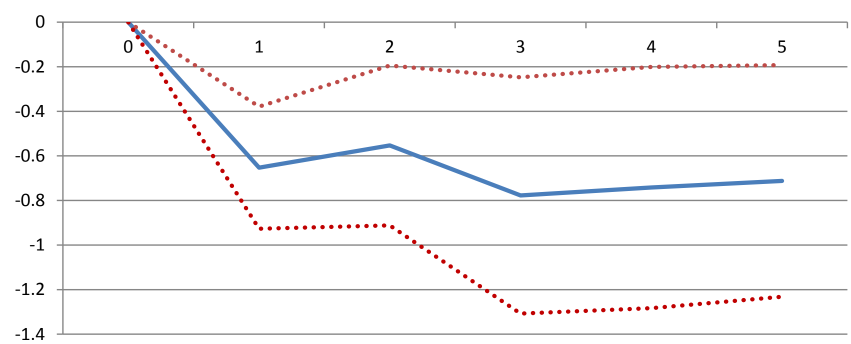
Note: IRFs are estimated using the specification in equation (4). The x-axis denotes time. t=0 is the year of the reform.

Fig. 16. The effect of capital account liberalization on poverty rates: the role of financial inclusion (percent)– Poverty headcount ratio at 3.10$ (2011 PPP).



Note: IRFs are estimated using the specification in equation (5). The x-axis denotes time. t=0 is the year of the reform.

Fig. 17. The effect of capital account liberalization on inequality, the role of financial crises, Gini (percent).



Note: IRFs are estimated using the specification in equation (6). The solid line corresponds to the IRF; dotted lines correspond to 90 percent confidence bands. The x-axis denotes time. t=0 is the year of the reform.

Fig. 18. The effect of capital account liberalization on the labor share, percentage points.

inclusion play a key role in shaping the distributional response to financial globalization. In particular, our results suggest that the impact on inequality tends to be significantly smaller in countries with strong levels of financial development and financial inclusion, and when they are not followed by episodes of financial crises. These results suggest that benefit-to-cost ratio of liberalization is higher past certain thresholds of financial development and inclusion. Economic policies designed to foster these necessary supporting conditions are beneficial in themselves and also help to mitigate adverse distributional consequences of financial integration.

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Appendix. Country coverage

Income Group	Country	Market Gini Range	Kaopen Range
High Income	Australia	1960–2010	1970–2010
	Austria	1981–2010	1970–2010
	Bahamas, The	1975–2004	1977–2010
	Barbados	1970–1997	1974–2010
	Belgium	1974–2010	1970–2010
	Canada	1963–2009	1970–2010

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(*continued*)

Income Group	Country	Market Gini Range	Kaopen Range
	Croatia	1986–2009	1996–2010
	Cyprus	1990–2009	1970–2010
	Czech Republic	1987–2010	1996–2010
	Denmark	1963–2010	1970–2010
	Estonia	1981–2010	1996–2010
	Finland	1971–2010	1970–2010
	France	1975–2010	1970–2010
	Germany	1967–2010	1970–2010
	Greece	1981–2010	1970–2010
	Hong Kong SAR, China	1973–2006	1970–2010
	Hungary	1962–2010	1986–2010
	Iceland	1992–2010	1970–2010
	Ireland	1976–2009	1970–2010
	Israel	1976–2005	1970–2010
	Italy	1967–2010	1970–2010
	Japan	1961–2010	1970–2010
	Korea, Rep.	1963–2010	1970–2010
	Malta	2000–2010	1972–2010
	Netherlands	1973–2010	1981–2010
	New Zealand	1963–2007	1970–2010
	Norway	1973–2010	1970–2010
	Poland	1970–2010	1986–2010
	Portugal	1980–2010	1970–2010
	Singapore	1972–2009	1970–2010
	Slovak Republic	1987–2010	1996–2010
	Slovenia	1987–2010	1996–2010
	Spain	1975–2010	1970–2010
	Sweden	1960–2010	1970–2010
	Switzerland	1980–2009	1996–2010
	Trinidad and Tobago	1976–2005	1970–2010
	United Kingdom	1960–2010	1970–2010
	United States	1960–2010	1970–2010
Middle Income	Albania	1996–2005	1995–2010
	Algeria	1986–2005	1970–2010
	Angola	1995–2005	1993–2010
	Argentina	1974–2010	1970–2010
	Armenia	1986–2007	1996–2010
	Azerbaijan	1976–2008	1996–2010
	Belarus	1981–2007	1996–2010
	Belize	1993–1999	1985–2010
	Bhutan	2003–2005	1985–2010
	Bolivia	1989–2007	1970–2010
	Bosnia and Herzegovina	1991–2005	1999–2010
	Botswana	1985–2005	1972–2010
	Brazil	1976–2009	1970–2010
	Bulgaria	1962–2010	1994–2010
	Cameroon	1983–2002	1970–2010
	Cape Verde	1989–2005	1982–2010
	Chile	1977–2009	1970–2010
	China	1974–2005	1984–2010
	Colombia	1962–2009	1970–2010
	Congo, Rep.	2005–2006	1970–2010
	Costa Rica	1977–2009	1970–2010
	Cote d'Ivoire	1978–2002	1970–2010
	Djibouti	1995–2005	1982–2010
	Dominican Republic	1986–2009	1970–2010
	Ecuador	1987–2009	1970–2010
	Egypt, Arab Rep.	1964–2008	1970–2010
	El Salvador	1965–2008	1970–2010
	Fiji	1977–1992	1975–2010
	Gabon	1975–1977	1970–2010
	Georgia	1981–2006	1996–2010
	Ghana	1987–2006	1970–2010
	Guatemala	1979–2006	1970–2010
	Guyana	1992–1999	1970–2010
	Honduras	1989–2009	1970–2010
	India	1960–2005	1970–2010
	Indonesia	1970–2010	1970–2010
	Iran, Islamic Rep.	1979–2005	1970–2010
	Iraq	2003–2004	1970–1994
	Jamaica	1968–2004	1970–2010
	Jordan	1973–2006	1970–2010
	Kazakhstan	1981–2006	1996–2010
	Lao PDR	1992–2007	1981–2010
	Latvia	1981–2010	1996–2010
	Lebanon	1997–2005	1970–2010

(*continued on next page*)

(*continued*)

Income Group	Country	Market Gini Range	Kaopen Range
	Lesotho	1986–2005	1972–2010
	Lithuania	1981–2010	1996–2010
	Macedonia, FYR	1989–2007	1997–2010
	Malaysia	1970–2005	1970–2010
	Mauritius	1972–2006	1972–2010
	Mexico	1968–2010	1970–2010
	Moldova	1981–2010	1996–2010
	Mongolia	1995–2006	1995–2010
	Morocco	1975–2007	1970–2010
	Namibia	1993–2005	1994–2010
	Nicaragua	1992–2005	1970–2010
	Nigeria	1980–2004	1970–2010
	Pakistan	1963–2005	1970–2010
	Panama	1974–2009	1970–2010
	Papua New Guinea	1995–2005	1979–2010
	Paraguay	1990–2009	1970–2010
	Peru	1981–2009	1970–2010
	Philippines	1963–2009	1970–2010
	Romania	1989–2010	1976–2010
	Russian Federation	1981–2009	1996–2010
	Senegal	1991–2005	1970–2010
	South Africa	1972–2005	1970–2010
	Sri Lanka	1979–2002	1970–2010
	St. Lucia	1995–2005	1983–2010
	Sudan	1968–1969	1970–2007
	Suriname	1999–2005	1982–2010
	Swaziland	1994–2005	1973–2010
	Thailand	1974–2004	1970–2010
	Tunisia	1974–2005	1970–2010
	Turkey	1977–2009	1970–2010
	Turkmenistan	1981–2005	1996–2010
	Ukraine	1976–2007	1996–2010
	Uruguay	1980–2009	1970–2010
	Uzbekistan	1981–2005	1996–2010
	Venezuela, RB	1967–2010	1970–2010
	Vietnam	1992–2006	1980–2010
	Yemen, Rep.	1992–2005	2002–2010
	Zambia	1967–2005	1970–2010
Low Income	Bangladesh	1973–2010	1976–2010
	Benin	2003–2006	1979–2010
	Burkina Faso	1994–2003	1988–2010
	Burundi	1992–2006	1970–2010
	Cambodia	1994–2004	1995–2010
	Central African Republic	1992–2003	1970–2010
	Chad	2002–2005	1970–2010
	Comoros	2002–2005	1981–2010
	Congo, Dem. Rep.	2005–2006	1970–2000
	Ethiopia	1980–2005	1970–2010
	Gambia, The	1992–2003	1971–2010
	Guinea	1991–2006	1970–2010
	Guinea-Bissau	1991–2005	1981–2010
	Haiti	1987–2001	1984–2010
	Kenya	1969–2005	1970–2010
	Kyrgyz Republic	1981–2007	1997–2010
	Liberia	2005–2007	1970–2010
	Madagascar	1980–2005	1970–2010
	Malawi	1977–2005	1970–2010
	Mali	1989–2006	1970–2010
	Mauritania	1987–2000	1970–2010
	Mozambique	1996–2005	1988–2010
	Nepal	1976–2004	1970–2010
	Niger	1992–2005	1970–2010
	Rwanda	1985–2006	1970–2010
	Sierra Leone	1976–2005	1970–2010
	Tajikistan	1981–2004	1997–2010
	Tanzania	1977–2001	1970–2010
	Togo	2005–2006	1970–2010
	Uganda	1989–2006	1970–2010
	Zimbabwe	1990–1995	1984–2010

Source: Solt (2009) and Chinn-Ito. Income groups based on World Bank classification.

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