

Income inequality: the aftermath of stock market liberalization in emerging markets

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Abstract

Early research has documented that the large-scale equity market liberalizations of the last decade led the subsequent rise in aggregate equity indices, investment booms, capital flows and economic growth. An important and unaddressed issue is the normative question of whether and how these reforms shifted the distribution of incomes in the aftermath of equity market liberalization. In careful empirical analysis, we find a pattern indicating that income share growth accrued almost wholly to the top quintile of the income distribution at the expense of a “middle class” that we define as the three middle quintiles of the income distribution. A surprising finding is that the lowest income share remained effectively unchanged in the event of liberalization. These patterns are robust to the inclusion of a wide variety of controls for global shocks, country-specific factors, and contemporaneously implemented privatization and stabilization policies.

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1. Introduction

In the latter half of the 1980s and the early years of the 1990s, the governments of over two dozen sovereign nations began to implement a wave of major economic reforms, which included capital account liberalization, privatization and/or a host of stabilization policies. This large-scale experiment has fuelled an active academic and popular debate on the causes and consequences of these reforms, in part because the aftermath of such reforms is important to nations which, to date, are considering liberalizing reforms. Empirical evidence on the consequences of these reforms will therefore be an important

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tool in assessing how best to implement similar liberalization policies in reform-minded countries.

In this context, a nascent but important body of empirical research has begun to emerge, analyzing important questions such as the relation of capital account liberalization to (i) emerging market equity prices (e.g., Bekaert and Harvey, 2000a,b; Kim and Singal, 2000; Henry, 2000a; Froot et al., 2001), (ii) liquidity (Levine and Zervos, 1998), (iii) private investment (Levine and Zervos, 1998; Henry, 2000b; Bekaert and Harvey, 2000a,b), (iv) equity flows (Bekaert et al., 2002) (v) and economic growth (Bekaert et al., 2001). This literature has found strong evidence that suggests capital account liberalization is associated with higher equity prices, lower cost of capital, investment booms, greater capital flows and higher growth.

To date, an important but unaddressed question in this literature is the issue of whether and how these reforms have shifted the *distribution* of incomes in the reforming countries. This is an important area of research for several reasons. For example, the finding in Bekaert et al. (2001) that average economic growth increased after liberalization raises normative issues about the allocation of the generated wealth. One would presumably evaluate the success of liberalizing reforms differently when average growth uniformly raised incomes for all quantiles of the distribution, from a finding that the average growth post-liberalization was only influenced by gains to the upper tails of the income distribution. For many countries that are still considering capital account liberalization, it is important to evaluate the benefits of market liberalization such as investment booms relative to the potential downsides of such policy reforms, so that future reforms may be tailored to alleviate any negative fallout from undertaking such economic reforms.

This paper takes a step in this direction by presenting the first set of results on the association of capital account liberalization with income inequality. In this paper, we heuristically and empirically describe the dynamics of the shifts in income distributions in a sample of 11 countries that undertook extensive economic reforms between 1986 and 1995¹. Prior to our discussion of the empirical work, we consider various mechanisms that link capital account liberalization to income inequality. We do not intend these mechanisms to be causal in either direction, but to provide a framework in which the finding of an empirical link between the two variables should not necessarily be discarded as spurious.

We analyze income distribution changes by comparing the size of three income shares before and after liberalization, conditional on a set of country-specific factors, and contemporaneous global shocks. We study the share of GDP held by the top quintile, the lowest quintile, and a group we will henceforth refer to as the *middle class*, which represents the sum of the three middle quintiles. Because we track the changes in income shares which plausibly respond slowly to economic reforms (relative to equity prices or the dividend yield), we analyze movements in the distribution over the “short run” which we take as the first 2–4 years beyond the year of the first liberalizing reform. This aspect of the analysis is discussed in more detail when we present our methodological framework and the construction of event windows. We focus mainly on capital account liberalization, but also control for other reforms such as privatization and stabilization policies that were

¹ Data for the remaining countries are noisy and sparse. Details of the data are given in Section 2 of the paper.

contemporaneously implemented. We will use the terms financial market liberalization, stock market liberalization and capital account liberalization synonymously for our purposes. Regression results in this study are given both for simple linear models as well as event-study models, which are somewhat incongruous in this context but useful nevertheless.

The principal findings of this research indicate systematic shifts in the income distribution in the first 4 years after a country's first liberalization reform. First, the raw data indicate that in the 11 emerging markets we consider, nine experienced a growth in the top quintile's share of income, and the mean share held by the top quintile rose by 1.3 percentage points. Second, one mildly surprising finding is that there is no discernible change in the mean income share of the lowest quintile. In the sample of 11 countries, four experience declines in the lowest quintile's income share, seven experience growth, and the mean share held by this group rises by 0.1%, which is statistically indiscernible from zero in regression analysis. This finding supports a recent finding in Dollar and Kraay (2001) that income growth of the poor is one-for-one with mean income growth of the population, which implies a constant income share for the poor. Based on these two strands of evidence and a basic adding-up theorem, there must be a reduction in the aggregate income share held by the middle class in the aftermath of capital account liberalization, and we find this to be the case. Specifically, in 9 of the 11 emerging markets in the sample, the aggregate income share of the middle class falls after liberalization. The raw data indicate a mean reduction of 1.45 percentage points in their share of income.

We test this pattern in careful regression analysis, to control for variations in domestic fundamentals, contemporaneous domestic and global business cycle effects, and country-specific effects. The regression results indicate that the income share of the middle class is strongly negatively associated with liberalization, and this relation persists in the presence of a wide variety of controls. In general, this finding is strongest for the 3rd quintile but holds statistically for the 2nd and 4th quintiles as well. In contrast, there is no statistically significant relation between liberalization and the income share of the lowest quintile in the presence of a rich set of controls. The empirical evidence for the upper quintile points to a positive and statistically significant relation between its income share and the event of liberalization, in accordance with the raw mean increases in their income shares. However, for certain regression specifications this relation attenuates. Because our regression estimates are sensitive to the choice of the regressor vector, we report a variety of results. We do not lend any causality interpretation to our results because of the well-known problems of identification. Instead, all of our results are presented as conditional correlations, and are open to varying interpretations, some of which we will discuss in the context of the regression results. It is important to stress that the described patterns hold for income shares. In the sample of emerging markets we consider, average income of the top quintile rose in all 11 nations, average income of the middle class as well as the lowest quintile rose in 10 of 11 nations.²

² The only exception is Nigeria, where the upper quintile's share of income after liberalization rose by 6.6 percentage points (a 13.6% increase), which is significantly higher than the mean increase of 1.3 percentage points in the upper quintile of all nations in the sample.

The key results obtained in this research complement, and should be viewed in the context of, some important research that precedes this, e.g., [Levine and Zervos \(1998\)](#), [Henry \(2000a,b\)](#), [Bekaert and Harvey \(2000a,b\)](#) and [Bekaert et al. \(2001\)](#). For instance, our finding that average middle class income share decreased while the average income share of the upper quintile simultaneously increased, coupled with the finding of [Bekaert et al. \(2000\)](#) that mean growth increased post-liberalization, indicates that the “pie grew” upon liberalization, but that the generated wealth was disproportionately allocated to the upper tail, suggesting a mean-increasing distributional shift. Without the result of [Bekaert et al.](#), our finding could not reject the hypothesis that liberalization leads greater inequality with a mean-reducing distributional shift. The entire body of results jointly raises normative issues about capital account liberalization and the subsequent welfare of the population, and questions of the trade-off in mean income growth and the concomitant growth in inequality. Although the latter topics are of academic interest, they are not explicitly discussed here, but will be addressed in future research.

To better comprehend the dynamics of these shifts in the distribution post-liberalization, we complement the regression analysis with nonparametric tests of stochastic dominance of the pre-liberalization and post-liberalization income distributions. Logically, combining our results with that of [Bekaert et al. \(2001\)](#) yields a very precise set of null hypotheses. First, patterns of income levels described in the current research indicate that the average post-liberalization distribution of incomes should exhibit First Order stochastic dominance (FOSD) over the pre-liberalization distribution. Similarly, SOSD should be obtained. Second, the finding that liberalization on average raises mean income growth in [Bekaert et al. \(2000\)](#), coupled with our finding that this growth appears to be driven entirely by the upper quintile, indicates that we should not reject Lorenz dominance.³ We provide formal tests of each of these hypotheses to provide a larger context for the results of our regression analysis.

An important issue that often arises in studying emerging markets phenomena is that of dating a country's first liberalizing reform. In this respect, our research was aided greatly by [Bekaert and Harvey \(2000a,b\)](#), and [Henry \(1999, 2000a\)](#), whose work has compiled official liberalization dates on overlapping sets of the sample of 11 liberalizing countries we consider.⁴ Also useful to this paper is the work of [Deninger and Squire \(1996\)](#), which has assembled panel data on socioeconomic metrics for a large cross section of nations. Although the focus of this research is on the association of financial reforms with income

³ Lorenz dominance, or mean-normalized second order stochastic dominance (SOSD) of income distribution y_1 over income distribution y_2 implies that any social welfare function that is increasing and concave in income will record higher levels of welfare in y_1 than in y_2 ; see [Shorrocks \(1983\)](#) or [Litchfield \(1999\)](#).

⁴ As a referee has pointed out to us, the issue of dating liberalizing events is a somewhat unsettled matter in the literature. [Bekaert and Harvey \(2000a,b\)](#) and [Henry \(2000a\)](#) are some of the primary sources, and [Kim and Singal \(2000\)](#) is another. To limit the discrepancies that arise from using multiple sources, we have used data from [Bekaert and Harvey \(2000a,b\)](#) and supplemented these with data from [Henry \(2000a\)](#) for those nations not covered in the former. To verify the robustness of our results, we have re-estimated all models and reproduced the figures with data from [Kim and Singal \(2000\)](#). This analysis has revealed noticeable differences in only the individual country effects, and made the estimation less precise due to a much smaller set of nations in Kim/Singal (only 6 nations in their data set have corresponding income shares). The graphs, and alternative dates, are available upon request from us.

inequality, it is directly related to the vast literature in growth that has studied amongst other issues, growth and income inequality (some examples include Barro, 1990; Barro and Salai-Martin, 1995; Perotti, 1996; Banerjee and Duflo, 2000; Dollar and Kraay, 2001). This paper integrates some aspects of the growth research into studying emerging markets phenomena, by exploring income inequality in the aftermath of stock market liberalization.

We will focus on the share of income owned by the j th quintile as the dependent variable for the greater part of our empirical analysis (where $j=l$ represents the lowest, $j=h$ represents the highest quintile and, aggregating the three middle quintiles yields a “middle class” with $j=m$). One might consider studying the variation in the Gini coefficient of inequality to analyze income distributions pre- and post-liberalization; one drawback of this measure, however, is its inability to distinguish between distributions that are unequal in very different ways, e.g., thick-tailed at the upper end, versus thick-tailed at the lower. Furthermore, income shares are the relevant variable to analyze in studying relative gains, and in inferring the process of resource allocation in the liberalizing nation. Normative analyses about the welfare of the middle class whose income share falls in the event of capital account liberalization must therefore be tempered to allow for the possibility that the level of income in the middle class rises after liberalization. As indicated earlier, middle class income increased contemporaneously with the decrease in their share of national income in 10 of the 11 nations in our sample of liberalizing nations.

The remainder of the paper is organized as follows. In Section 2 we discuss the sample of emerging market nations in our study, provide some preliminary findings, and describe various mechanisms that link capital account liberalization with the observed changes in income shares. Section 3 describes our empirical methodology; results from both standard regression analysis and event study models are reported here. In Section 4, we discuss tests of stochastic dominance and Lorenz dominance, then describe the findings of our tests. Section 5 considers whether it is possible to give a causality interpretation to our findings. Conclusions follow. Appendix A details the data construction of this study, and the construction of some of our test statistics.

2. Capital account liberalization and income inequality

2.1. Sample of liberalizing and non-liberalizing nations

Listed in Table 1a, our panel data consist of 11 countries that underwent capital account liberalization between 1986 and 1995, and a sample of eight “control” countries that were subject to no major capital account reforms in this time period. The liberalizing countries are Brazil, India, Korea, Malaysia, Mexico, Nigeria, Pakistan, Philippines, Sri Lanka, Thailand and Turkey. Although over two-dozen sovereign nations implemented financial market liberalization in varying degrees in this time period, our sample is truncated for unavailability of income share data. For example, only pre-liberalization income share data are available for Chile and Morocco, while no income share data are available for other liberalizing countries such as Argentina, Taiwan and Venezuela. The control countries in our sample are Bangladesh, Cote d’Ivoire, Guatemala, Jamaica, Kenya, Niger, Sierra Leone and Trinidad and Tobago. The fact that these represent the same regions as the

Table 1a
Sample of liberalized and non-liberalized countries

Liberalized countries	Liberalization date	T_i	Non-liberalized countries	T_i
Brazil	1988	13	Bangladesh	11
India	1986	13	Cote d'Ivoire	6
Korea, Rep.	1987	6	Guatemala	3
Malaysia	1987	4	Jamaica	9
Mexico	1989	5	Kenya	2
Nigeria	1995	5	Niger	3
Pakistan	1991	8	Sierra Leone	2
Philippines	1986	6	Trinidad and Tobago	5
Sri Lanka	1992	7		
Thailand	1988	7		
Turkey	1989	3		

The financial liberalization dates are based on [Bekaert, Harvey and Lundblad \(2001\)](#) and [Henry \(2000a\)](#). T_i indicates the number of years for which income shares are available for each country in the sample. National-level data on income shares of the various quintiles of the population were obtained from the High Quality Sample of [Deininger and Squire \(1996\)](#) and from the [World Income Inequality Database \(2000\)](#).

The minimal criterion for including a country in the sample was the availability of a minimum of one data point in each of the 5-year windows prior to and subsequent to liberalization.

liberalizing countries mitigates possible region-specific effects when comparing distributions between the liberalizing and control countries.

Annual income share data are compiled both from [Deninger and Squire \(1996\)](#) and the World Income Inequality Database (WIID). These data, which are available for T_i years for county i , yield an unbalanced panel for our data set. All stock market data, deflated by the price index, are taken from World Bank Database on Financial Structure and Economic Development (FSED). Country-specific data are obtained from the World Bank's Socio-economic Time Series Access and Retrieval (STARS) database, and rule of law data are from the IRIS data set. A description of the variables is provided in Appendix A. For dates of first capital account liberalization, we lean on the careful dating procedures developed by [Bekaert and Harvey \(2000a,b\)](#), and [Henry \(1999, 2000a\)](#). The liberalization dates are presented in [Table 1a](#). Summary statistics of the key variables are in [Table 1b](#). All data are measured at the annual frequency.

2.2. Mechanisms

There are a few important channels by which stock market liberalization could differentially affect different income groups in the population. Standard international asset pricing models (e.g., [Stulz, 1995a,b](#)) predict that global capital market integration may reduce the domestic cost of equity capital.⁵ Although the domestic risk-free rate could rise above its autarky rate in the event of liberalization, research preceding ours indicates that the equity premium falls in the sample of 12 emerging markets we consider (see, for example, [Tesar and Werner, 1998](#); [Stulz, 1999](#); [Bekaert and Harvey, 2000a,b](#); [Henry,](#)

⁵ For instance, the presence of domestic investment opportunities could lead to net capital inflows that lower the riskless rate; and, greater risk sharing between domestic and foreign investors could result in reducing the risk premium (e.g., [Errunza and Losq, 1985](#)).

2000a). If equity market liberalization reduces the cost of equity capital, then holding future cash flows fixed, an immediate implication is an increase in the country's equity price index and, from a valuation perspective, capital gains in the liberalizing country at least over the short horizon. However, if participation in the stock market is segmented by income groups, then these gains will accrue only to those income groups that own equity, and *ceteris paribus* the measured response to liberalization will be higher for those groups whose income depends directly on the equity market. Although we have very sparse data on stock market participation in emerging markets, there is anecdotal evidence to strongly suggest that equity ownership is largely restricted to the upper income classes in the countries we sample (e.g., [Beim and Calomiris, 2000](#)).

Over the short horizon, stock market liberalization may also differentially impact the different income groups if access to credit markets is disparate within the population. By reducing the cost of capital, stock market liberalizations increase the net present valuation (NPV) of ongoing investments, while also rendering previous negative NPV projects feasible under a lowered cost of external finance.⁶ Since there is evidence that growth is strongly positively associated with lowered cost of financing (see [Rajan and Zingales, 1998](#)), there will be a measured increase in income via entrepreneurial wealth and retained earnings following the event of stock market liberalization. However, if credit market imperfections limit lower income groups' access to credit, as in the models of [Galor and Zeira \(1993\)](#) and [Ray \(1998\)](#), such gains will be disproportionately allocated to the groups with better access to bank credit and external financing.

Capital account liberalization may also lead to distributional changes by affecting factors unobservable to the analyst. For example, the implementation of stock market liberalization might signal a change in domestic fundamentals that leads firms to expect a better environment for growth, as hypothesized in [Bartolini and Drazen \(1997\)](#). Improved expectations might be particularly strong for firms that are closely tied to the domestic government because they are in the best position to appropriate the gains from capital market integration, as implied by the rent-seeking models of [Krueger \(1974\)](#) and [Bhagwati \(1996\)](#). In particular, favored firms or investors may receive better information, or explicitly influence reforms that reap private benefits. To the extent that members of the upper income quintiles are tightly linked with the domestic governments, these factors should most strongly impact their income shares post-liberalization. Therefore, an *a priori* assumption is that the measured income of the upper quantiles will be inflated via this channel. If these groups undertake investments because they have private information of their prospects, then one would observe an increase in their mean income share even after controlling for contemporaneous increases in stock market valuation and investment. One might speculate, in the spirit of the [Kuznets's \(1955\)](#) inverted-U hypothesis,⁷ that stock market liberalization leads to improvements in the upper tails of the distribution at the early stage of stock market development, while the lower quintiles gain over time.

The preceding discussion is intended to indicate possible links between liberalizing reforms and changes in income shares. We do not press on the issue of causality since

⁶ [Henry \(2000b\)](#) finds evidence that stock market liberalization leads investment booms.

⁷ Kuznet's hypothesis is an initial rise in inequality with economic progress and then a gradual fall as the benefits of growth permeate more widely.

Table 1b
Summary statistics for liberalizing countries

	Brazil	India	Korea	Malaysia	Mexico	Nigeria	Pakistan	Philippines	Sri Lanka	Thailand	Turkey
<i>GDP per capita, PPP (current international \$)</i>											
Pre-liberalization											
Mean	4565.88	783.92	4110.91	3380.06	5230.02	769.90	1167.33	2374.05	1904.86	2096.13	3606.15
Std. Deviation	798.89	73.88	815.25	286.48	450.75	48.79	151.76	105.13	140.49	372.67	642.95
Post-liberalization											
Mean	5669.59	1292.65	8779.73	5419.37	6858.15	806.87	1662.13	2957.79	2700.13	4250.25	5118.65
Std. Deviation	267.48	119.03	1229.37	673.62	481.50	10.44	91.66	175.61	213.44	657.95	416.53
<i>School enrollment, secondary (% gross)</i>											
Pre-Liberalization											
Mean	35.80	34.96	88.50	52.22	55.90	28.82	19.36	65.58	72.34	29.96	42.86
Std. Deviation	1.33	2.68	4.95	2.68	0.42	3.49	2.02	1.82	1.46	1.09	2.20
Post-Liberalization											
Mean	40.17	41.27	90.95	57.52	54.78	N.A	25.60	71.15	74.73	33.25	50.58
Std. Deviation	2.22	3.27	1.47	0.73	1.96	N.A	0.00	3.16	0.59	5.71	3.94
<i>General government consumption (% of GDP)</i>											
Pre-Liberalization											
Mean	10.13	10.23	10.83	16.21	8.95	15.48	14.75	8.16	9.97	12.73	7.77
Std. Deviation	1.43	0.54	0.70	1.43	0.35	2.51	1.60	0.85	0.29	0.84	0.33
Post-liberalization											
Mean	16.46	11.66	10.35	14.18	9.70	9.59	12.72	9.16	10.14	9.68	11.70
Std. Deviation	2.52	0.34	0.37	0.76	1.37	1.55	0.91	0.86	0.83	0.34	1.39
<i>Rule of law index</i>											
Pre-liberalization											
Mean	3.82	2.70	2.76	4.54	3.22	2.38	1.82	1.00	0.36	3.00	3.00
Std. Deviation	0.25	0.48	0.43	0.51	0.44	0.91	0.40	0.00	0.50	0.00	0.00

Post-liberalization											
Mean	3.97	1.73	2.63	3.50	3.00	3.00	2.30	1.00	2.90	4.20	3.18
Std. Deviation	0.08	0.42	1.20	0.84	0.00	0.00	1.03	0.00	1.39	0.44	1.01

Private credit by deposit money banks and other financial institutions/GDP

Pre-liberalization											
Mean	0.21	0.28	0.66	0.73	0.10	0.10	0.26	0.39	0.19	0.53	0.16
Std. Deviation	0.07	0.01	0.05	0.20	0.01	0.01	0.02	0.06	0.01	0.03	0.00
Post-liberalization											
Mean	0.25	0.28	0.86	0.81	0.23	0.08	0.23	0.19	0.22	0.77	0.13
Std. Deviation	0.06	0.03	0.11	0.11	0.09	0.00	0.01	0.02	0.02	0.16	0.00

Stock market capitalization/GDP

Pre-liberalization											
Mean	0.12	0.04	0.07	0.58	0.04	0.05	0.06	0.05	0.09	0.05	0.02
Std. Deviation	0.04	0.01	0.02	0.04	0.02	0.01	0.01	0.02	0.04	0.02	0.00
Post-liberalization											
Mean	0.11	0.10	0.37	0.91	0.26	0.08	0.18	0.13	0.17	0.35	0.10
Std. Deviation	0.05	0.04	0.13	0.32	0.16	0.02	0.04	0.06	0.04	0.22	0.05

Stock market value traded/GDP

Pre-liberalization											
Mean	0.06	0.02	0.05	0.07	0.04	0.00	0.01	0.01	0.00	0.03	0.00
Std. Deviation	0.04	0.01	0.03	0.03	0.04	0.00	0.00	0.00	0.00	0.03	0.00
Post-liberalization											
Mean	0.06	0.06	0.35	0.20	0.11	0.00	0.05	0.03	0.03	0.36	0.07
Std. Deviation	0.04	0.02	0.13	0.10	0.06	0.00	0.03	0.01	0.02	0.25	0.06

Summary statistics are for the 5 years prior to and the 5 years subsequent to liberalization.

liberalizations are rarely implemented in a vacuum, as discussed in the next section. Rather, the above discussion intends to convey the notion that were liberalizations implemented exogenously, there are mechanisms by which they may directly or indirectly affect income shares. We will investigate the empirical links between income shares and liberalization in the next section. Prior to discussing the methodological framework employed in the regression analysis, we present some preliminary findings.

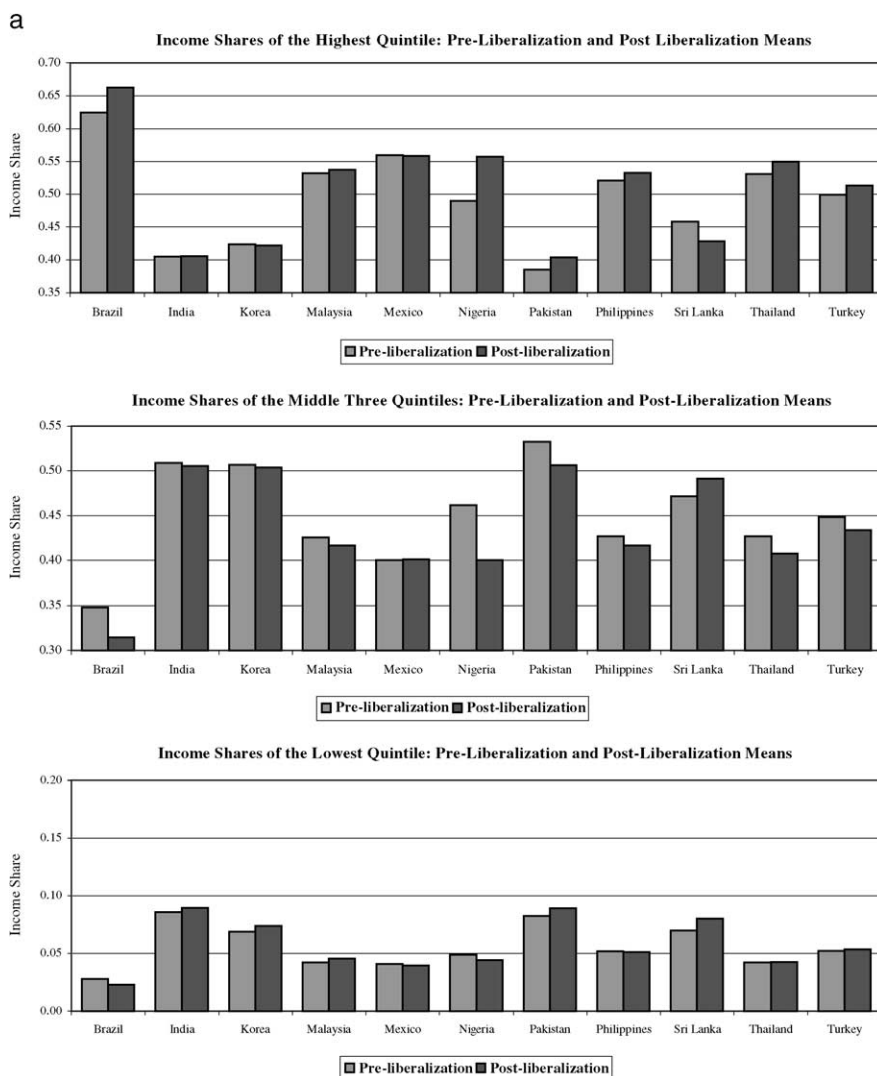


Fig. 1a. Income shares pre- and post-liberalization.

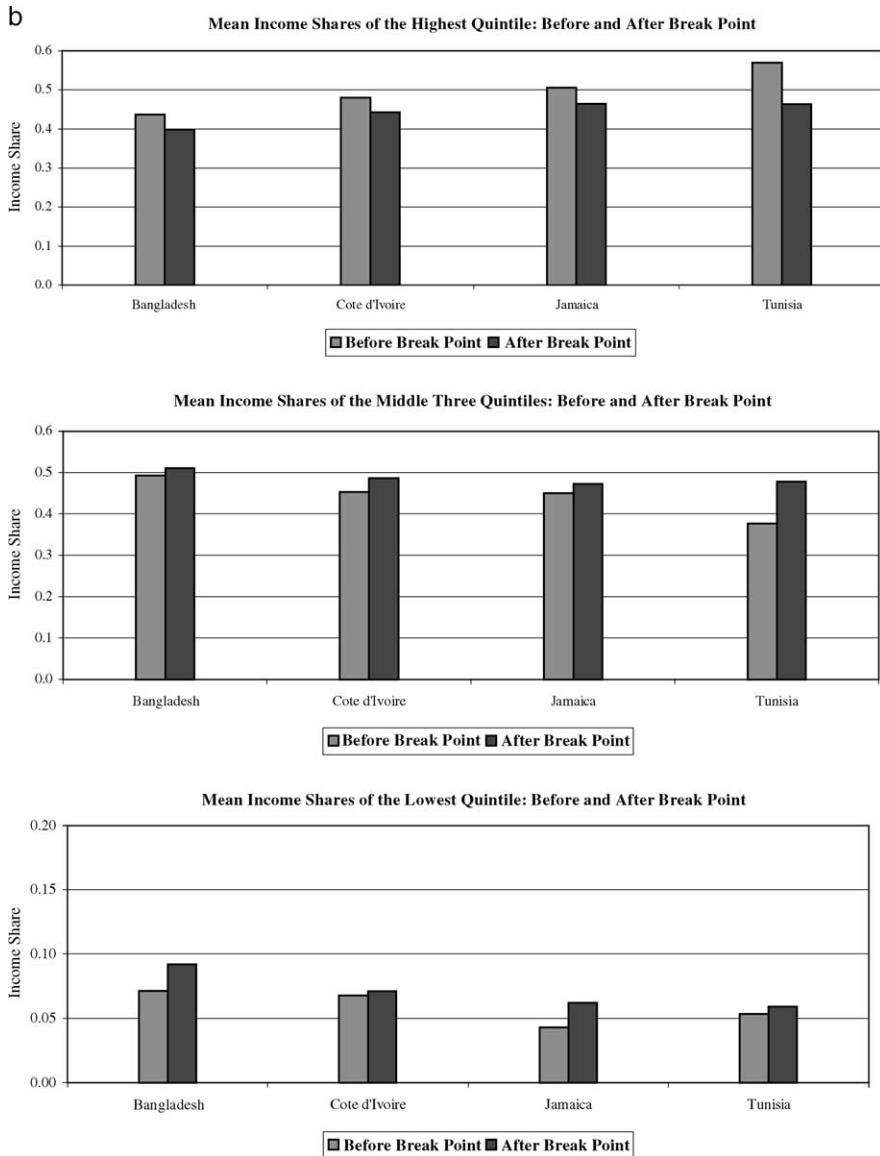


Fig. 1b. Income shares in non-liberalizing nations.

2.3. Preliminary findings

The central findings of this research are summarized in Fig. 1a. These graphs, which display mean income shares pre- and post-liberalization, reveal three main shifts in the

Table 2
Mean income shares and gini coefficients

Countries	Income share of top quintile		Income share of middle quintile		Income share of bottom quintile		Gini coefficient	
	Pre-liberalization	Post-liberalization	Pre-liberalization	Post-liberalization	Pre-liberalization	Post-liberalization	Pre-liberalization	Post-liberalization
Brazil	0.624	0.662	0.348	0.315	0.028	0.023	57.43	60.79
India	0.405	0.406	0.509	0.505	0.086	0.089	31.49	31.31
Korea	0.424	0.422	0.507	0.504	0.069	0.074	35.12	33.64
Malaysia	0.532	0.537	0.426	0.417	0.042	0.046	48.00	48.35
Mexico	0.559	0.559	0.400	0.401	0.041	0.040	50.58	51.34
Nigeria	0.490	0.557	0.462	0.400	0.049	0.044	42.46	50.60
Pakistan	0.385	0.405	0.532	0.506	0.082	0.089	32.17	31.19
Philippines	0.521	0.533	0.427	0.417	0.052	0.051	46.08	46.72
Sri Lanka	0.459	0.428	0.471	0.491	0.070	0.080	38.40	34.40
Thailand	0.531	0.550	0.427	0.407	0.042	0.043	47.40	47.22
Turkey	0.499	0.513	0.448	0.434	0.052	0.054	44.17	45.25

Four-year pre-liberalization and post-liberalization windows were used to calculate the income shares and gini coefficients.

For nations with fewer than 4 annual years of data, the maximum available years were used to construct the means.

income distribution in the first 4 years after a country's first liberalization reform.⁸ First, in almost all countries in the sample, income shares of the highest quintile increased in the 4 years post-liberalization. Second, there is a (weak) reduction in aggregate income shares of the three middle quintiles, in varying degrees, in the aftermath of stock market liberalization. Third, there is a mixed reaction for the lowest income group for our sample, as this share rose in some countries in the sample, and fell in other countries. In particular, the mean increase in the countries where the lowest income share rose is found to be statistically indistinguishable from the mean decrease in those countries where the lowest income share fell, making it difficult to detect any meaningful shift in their income share in the event of liberalization.

As a comparison, we also plot the income shares in non-liberalizing countries in Fig. 1b. Because there is no liberalization date, we pick the break point to be 1989 (which is the median liberalizing year in our sample) and compare income shares before and after 1989.⁹ The idea of using a common break point has previously been studied by Levine and Zervos (1998). As Fig. 1b indicates, there is no striking pattern that emerges in income shares of non-liberalizing nations in this time period. Highest income quintiles both rise and fall after the breakpoint, and similarly for the middle income quintiles. The lowest quintile shares uniformly rise after the breakpoint.

The fact that income shares of the top quintiles grow, while those of the remaining quintiles shrink or remain stable, indicates a growth in inequality in the nations studied. Another way to examine whether inequality grew after liberalization is to examine the size of the well-known Gini coefficient of inequality before and after liberalization. This is done in Table 2 and the results corroborate our hypothesis.

The preliminary finding, based on Fig. 1a,b and Table 2, indicates that the income distribution is altered in the event of a liberalizing reform. We investigate this pattern more thoroughly in the context of regression analysis in the next section.

3. Empirical methodology

Two empirical strategies are used to study the empirical links between capital account liberalization and income inequality. In this section we attempt to uncover the relation between liberalization and income shares using regression analysis, and an event-study model that permits a diffused effect of liberalization on income shares. In the following section, we verify the changes in income distribution implied by the regression results, by tests of stochastic dominance and Lorenz dominance of the income distribution.

⁸ Where possible the means are computed over the 4 years preceding and 4 years following the first year of liberalization. For those nations with fewer than 4 years of data before and after liberalization (see Table 1a), individual means are computed over as many observations available.

⁹ We also tested this with alternate break points, with no substantive differences in the results to report.

3.1. Income shares and stock market liberalization

As indicated in the discussion of the mechanisms that link capital account liberalization to income inequality, valuation and capitalization changes in the local equity market after a liberalizing reform may be an immediate link between income shares and liberalization. Therefore, controlling for variations in equity markets across emerging markets, and over time within them, is expedient. We focus on two measures of equity market activity, which have been previously been considered (e.g., [Atje and Jovanovic, 1993](#), [Levine and Zervos, 1998](#); [Bekaert et al., 2001](#)). For an indicator of the activity in a local equity market, we use the actual dollar valuation of stock traded normalized by the local GDP. We also control for the size (or, financial sector development) of the local market using the equity market capitalization normalized by GDP.

The regression framework explicitly accommodates a delayed effect of stock market variables on income shares. As in [Barro \(1990\)](#) and [Blanchard et al. \(1993\)](#), both contemporaneous and lagged values of stock market and income variables will be utilized. Strictly speaking, the association of year t revaluations of the equity market with year t income shares is not “contemporaneous” in these data, because these data are annual averages.¹⁰

The first set of conditional correlations are obtained from estimating variants of the regression model:

$$Q_{jit} = \alpha_i + \tau_t + y \times \text{PostLiberalize}_{it} + X'_{it} \beta + \delta_1 \times \text{SMC}_{it} + \delta_2 \times \text{SMC}_{it-1} + \delta_3 \\ \times \text{SMV}_{it} + \delta_4 \times \text{SMV}_{it-1} + \varepsilon_{it},$$

where

$$\text{PostLiberalize}_{it} = \begin{cases} 1 & \text{if year } t \text{ is year of or any year after } i\text{'s first capital account liberalization policy} \\ 0 & \text{otherwise} \end{cases}$$

i indexes the country, t indexes the year, SMC represents our measure of equity market size and SMV (an acronym for stock market valuation) will denote our measure of equity market activity described earlier. X is a vector of country-specific controls that include the contemporaneous and lagged values of the logarithm of per capita GDP, secondary school enrollment, government consumption, gross investment, and one measure of the “rule of law” as in [Alesina and Rodrik \(1995\)](#).¹¹ To partially capture differences in credit markets, a measure of banking sector development (measured by aggregate private credit normalized by GDP) is added as well. The liberalization indicator is included in addition to the year dummies to isolate the effects of liberalization from those of contemporaneous global or regional shocks.

¹⁰ This accommodates the possibility that changes in the equity market earlier in the year are associated with income changes later in the year.

¹¹ The construction of this variable is detailed in Appendix A.

Table 3a
Changes in income shares around financial market liberalization

	(A)			(B)		
	Q_h	Q_m	Q_l	Q_h	Q_m	Q_l
	(1)	(2)	(3)	(4)	(5)	(6)
PostLiberalize _{<i>t</i>}	0.036** (0.013)	− 0.031** (0.011)	− 0.0053* (0.003)	0.057** (0.0259)	− 0.048** (0.019)	− 0.009 (0.007)
LnGDP _{<i>t</i>}				0.206* (0.177)	− 0.210* (0.117)	− 0.006 (0.084)
LnGDP _{<i>t-1</i>}				0.061 (0.236)	− 0.082 (0.371)	0.020 (0.072)
SMC _{<i>t</i>}				− 0.293** (0.146)	0.247** (0.109)	0.045 (0.044)
SMC _{<i>t-1</i>}				0.269 (0.282)	− 0.184 (0.210)	− 0.086 (0.086)
SMV _{<i>t</i>}				0.243* (0.141)	− 0.216* (0.148)	− 0.023 (0.061)
SMV _{<i>t-1</i>}				0.057 (0.498)	− 0.081 (0.176)	0.025 (0.151)
GovtConsumption				− 0.001 (0.005)	0.001 (0.004)	0.001 (0.001)
Secondary Enrollment				− 0.002 (0.002)	0.001 (0.001)	0.001 (0.002)
PrivateCredit				− 0.077 (0.182)	0.060 (0.136)	0.015 (0.055)
Rule of Law				0.082* (0.051)	− 0.070* (0.038)	− 0.011 (0.015)
Terms of Trade				0.169 (0.155)	0.111 (0.115)	0.569 (0.472)
Year Dummies	✓	✓	✓	✓	✓	✓
Country Fixed Effects	✓	✓	✓	✓	✓	✓
N	113	113	113	80	80	80

The dependent variable is the income share of the Q_j^{th} income group, where Q_h , Q_m and Q_l are defined in the text. The explanatory variables are: *PostLiberalize* for an indicator for liberalization, *LnGDP_t* and *LnGDP_{t-1}* for the current and lagged values of GDP respectively; *SMC_t* and *SMC_{t-1}* for the current and lagged values of normalized stock market capitalization respectively, and *SMV_t* and *SMV_{t-1}* for the current and lagged values of normalized stock market valuation respectively. The explanatory variables also include a measure of the banking sector development *Private Credit*, measures of secondary schooling, Rule of law, terms of trade, as well as country-specific dummies, and year dummies. The reported standard errors are heteroskedasticity-consistent (White) and reported in parentheses. * and ** respectively denote significance at the 10 and 5 percent error levels.

Panel (A) in Table 3a reports regressions of Q_j ($j=l, m, h$) on the liberalization indicator and a full set of year dummies and country-specific effects, yielding conditional mean differences in income shares pre- and post-liberalization. The results show a strong positive (respectively, negative) relationship between liberalization and Q_h (respectively, Q_m). In Column (1) the coefficient on *PostLiberalize* implies a mean increase of 3.6 percentage points in the income share of the highest quintile after liberalization. Column (2) indicates a mean decrease of 3.1 percentage points in the Q_m regression. In Column (3) the coefficient is significant, but at a much higher error level, pointing to a 0.5 percentage point increase in the income share of the lowest quintile. Together these results

corroborate the pattern found in Fig. 1. However, these differences are larger in absolute value than the simple averages of income shares before and after liberalization, indicating that unobservable country-specific factors and contemporaneous global or regional shocks may be correlated with the decision to liberalize. This issue is further researched below where the regression model controls for coincidental implementation of major domestic reforms.

Panel (B) of Table 3a presents results when the above model is supplemented with a wide variety of controls for variations in equity market, banking sector and domestic fundamentals. These results show that while the basic pattern that emerged in Panel (A) persists in the presence of the additional covariates, the magnitude of that pattern is inflated with the additional controls, as evinced by the absolute value on PostLiberalize_t . Furthermore, the coefficient on PostLiberalize is no longer significant in Column (6). In Column (5), controlling for equity market size via SMC_t , the coefficient on SMV_t is found to be negative and significant, suggesting that equity market activity is inversely associated with the middle class's share of income. Concurrently, the coefficient on SMV_t is positive in Column (4), suggesting that the upper quintile's income share is affected by contemporaneous changes in equity revaluation. One possible mechanism for this finding is that greater equity market activity could plausibly reflect the greater willingness of participants to hold assets, and this leads to equity price appreciation, benefiting the upper quintile (who are the typical owners of equity in the sample we consider). The results reveal no statistical association between the lowest income quintile and equity market activity; correspondingly, the point-estimates on SMV_t in the Q_h and Q_m regressions approximately sum to zero. Neither do the results reveal any statistical association between lagged values of equity market measures and income shares. Since liberalizations are associated with higher aggregate stock prices (Henry (2000a)), these results are consistent with a scenario in which the upper income quintile disproportionately participates in the equity market, accruing the capital gains from equity revaluation.

Some discussion of the additional regressors is warranted. We note that the measure of the banking sector development, PrivateCredit , is statistically insignificant in all specifications. Nevertheless, the central tendency of the estimator is positive for each of the Q_m and Q_l regressions indicating, perhaps, that banking sector development leads to better access for lower income groups, as discussed in Section 2.2. In a related vein, the coefficient on SMC_t is worth further analysis. Suppose that growth in the financial sector is associated with greater stock market participation. Although we do not have data on participation, there is some casual evidence that participation is largely restricted to the upper quintiles in emerging markets (see Beim and Calomiris, 2000). Growth in the financial sector might therefore be associated with a greater share of the middle class' participation in the equity market, which would divert some fraction of the short-run capital gains from liberalization to the middle class. In this scenario, we would observe both, a positive coefficient on SMC_t for the middle class, and a negative coefficient for the upper quintile. This is found to be the case in Columns (5) and (4), respectively.

In Table 3b we consider a small extension of the previous results by investigating how the income distribution is affected by differences in the strength of legal institutions/government stability, by introducing the interaction $\text{Rule of Law} \times \text{PostLiberalize}$. Essen-

Table 3b

Rule of law and liberation: different effects

	Q_h	Q_m	Q_l
	(1)	(2)	(3)
PostLiberalize	0.049** (0.0259)	– 0.038** (0.017)	– 0.007 (0.007)
Rule of Law	0.084* (0.051)	– 0.076* (0.038)	– 0.013 (0.015)
PostLiberalize $\times$ Rule of Law	0.009* (0.005)	0.013 (0.142)	0.010 (0.221)
Year Dummies	✓	✓	✓
Country Fixed Effects	✓	✓	✓

This table reports results for a regression that replicates the regression model in Table 3a, including all the regressors from Table 3a, plus an interaction between Rule of Law_{*t*} and PostLiberalize_{*t*}. The reported standard errors are heteroskedasticity-consistent (White) and reported in parentheses. * and ** respectively denote significance at the 10% and 5% error levels.

tially, the hypothesis is that although stock market liberalization has been found to positively affect income shares, the efficacy of such liberalizing policies might depend on the perceived and actual stability of the government and other legal institutions. In Table 3b we find that this is indeed the case, although the hypothesis is statistically verified only for the top quintile. Specifically, we find that while a generic discrete jump in Rule of Law from 0 to 1 (i.e., anarchy to fully accountable) would result in a 8.2% appreciation in the top quintile's income share, a liberalized-induced jump in Rule Of Law would result in a larger 9.3% appreciation of the top quintile's income share. However, we are unable to statistically discern such a pattern for any of the other income shares.

3.2. Liberalization-induced changes in income shares

The results in Table 3a indicate that equity market revaluations affect income shares independently of liberalizing reforms, since the liberalizing indicator is a control variable in the estimated regression model. A legitimate question to consider is whether the association of income shares and equity market activity is systematically different prior and subsequent to liberalization. It is plausible that liberalization-induced revaluations in equity markets elicit different responses from investors than generic changes in equity market revaluations. This is plausible because the event of a liberalizing reform signals fundamental changes that lead firms to expect a better environment for growth, as discussed in Bartolini and Drazen (1997). For example, Henry (2000a) finds that private investment is differentially associated with liberalization-specific and generic changes in the rate of return.

To test this hypothesis, the above regression model was modified as follows:

$$\begin{aligned}
 Q_{jit} = & \alpha_i + \tau_t + \gamma \text{PostLiberalize}_{it} + X'_{it} \beta + \delta_1 \text{SMC}_{it} + \delta_2 \times \text{SMC}_{it-1} + \delta_3 \times \text{SMV}_{it} \\
 & + \delta_4 \times \text{SMV}_{it-1} + \delta_5 \times \text{SMV}_{it} \times \text{PostLiberalize}_{it} + \delta_6 \times \text{SMV}_{it-1} \\
 & \times \text{PostLiberalize}_{it} + \varepsilon_{it}.
 \end{aligned}$$

The inclusion of the interactive variables permits a differential slope effect of equity market activity prior and subsequent to a nation's liberalizing reform. In particular, if

liberalization-induced changes in stock market activity have fundamentally different implications for incomes from generic changes in stock market valuation, then we must find that $\delta_5 \neq 0$, and $\delta_6 \neq 0$. Estimated results are presented in Table 4.

The coefficients on both contemporaneous SMV_t , as well as its interaction with the liberalization indicator, are both statistically significant at the 10% error level, but of differing signs. In the Q_m regression, the coefficients δ_3 and δ_5 jointly indicate that while a generic 10% increase in equity market activity leads to a 2.8 percentage point decrease in Q_m , a liberalization-induced change of the same magnitude leads to a somewhat larger decrease of 3.5 percentage points. The results reject neither the hypothesis that $\delta_4=0$ nor $\delta_6=0$, indicating that changes in equity market revaluation in the preceding year have no statistically significant effect on Q_m . Similarly, the results indicate that a 10% increase in equity market activity tends to have a liberalization-induced increase in Q_h that is 2.6 percentage points higher than that for a generic 1% increase in trading activity. As with Q_m , there is no observed statistical association between the lagged measure of stock market activity and changes in the upper quintile's income share. The data also do not support any effects of changes in equity market activity before or after liberalization on the lowest income share.

A model of differential access to credit markets across income groups may reconcile the liberalization-induced slope effects found in Table 4. Since investment is (i) spurred after liberalization and (ii) more sensitive to liberalization-specific changes rather than generic changes in valuations (Henry (2000b)), liberalizing reforms are likely to lead to greater entrepreneurial wealth and retained earnings via the investment channel. However, as discussed earlier, segmented access to credit markets may result in these gains accumulating disproportionately to higher income groups. The differential slopes are also consistent with a story of the improved expectations after liberalizing reforms due, for example, to higher risk-sharing (see Bekaert and Harvey, 1995; Bartolini and Drazen, 1997). If liberalization does portend a differential association of changes in equity market valuations with income shares, then a differential association must also emerge between liberalizing nations and non-liberalizing nations. We test this hypothesis next.

Table 4
Liberalization-induced versus generic changes in stock market valuation

	Q_h	Q_m	Q_l
PostLiberalize	0.054* (0.028)	−0.045** (0.020)	−0.004 (0.006)
SMV_t	−0.349** (0.165)	0.281** (0.122)	0.066 (0.052)
SMV_{t-1}	0.356 (0.318)	−0.244 (0.233)	−0.111 (0.100)
PostLiberalize $\times$ SMV_t	0.348* (0.213)	−0.270* (0.163)	−0.007 (0.046)
PostLiberalize $\times$ SMV_{t-1}	−0.299 (0.332)	0.220 (0.245)	−0.010 (0.069)
Year Effects	✓	✓	✓
Country Dummies	✓	✓	✓

The dependent variable is the income share of the Q_j^{th} income group ($j=h,m,l$). The explanatory variables include PostLiberalize* SMV_t which the interaction of PostLiberalize and SMV_t and PostLiberalize* SMV_{t-1} which is similarly defined as the interaction of its constituent variables. Country-specific dummies, year dummies, and all the control variables from Table 3a were estimated but are not reported for brevity. The reported standard errors are heteroskedasticity-consistent (White) and reported in parentheses. * and ** denote significance at the 10 and 5 percent error level, respectively.

3.3. Income shares in non-liberalization nations

A natural concern is that the results obtained in [Tables 3a and 4](#) are driven, not by liberalizing reforms in the emerging markets, but rather by world or regional business cycle effects that are not adequately captured by year dummies.¹² If global economic changes are contemporaneously affecting all nations (or regional ones are affecting all nations in the particular region), including those that did not implement any reforms in this time-period, it may be difficult to conclude that liberalization has any statistical effect on income shares. Here we conduct a simple regression test to determine the validity of this alternative explanation.

$$\begin{aligned} Q_{jit} = & \alpha_i + \tau_t + \gamma_1 \times \text{PostLiberalize}_{it} + \gamma_2 \times \text{NonLiberalize}_i + X'_{it} \beta + \delta_1 \times \text{SMC}_{it} \\ & + \delta_2 \times \text{SMC}_{it-1} + \delta_3 \times \text{SMV}_{it} + \delta_4 \times \text{SMV}_{it-1} + \delta_5 \times \text{SMV}_{it} \\ & \times \text{NonLiberalize}_i + \delta_6 \times \text{SMV}_{it-1} \times \text{NonLiberalize}_i + \varepsilon_{it}. \end{aligned}$$

where

$$\text{NonLiberalize}_i = \begin{cases} 1 & \text{if country } i \text{ has not undertaken any liberalization reform} \\ 0 & \text{otherwise.} \end{cases}$$

The indicator NonLiberalize, which does not vary over time in our sample, takes on the value 1 for the set of “control” nations listed in [Table 1a](#). These controls should help mitigate the presence of unobservable regional differences between the controls and the liberalizing nations because both sets of countries represent the same world regions.

The results for variations of this model are given in [Table 5](#). In Panel (A) we report results for a regression model with no interaction variables. In Column (1), the coefficient on NonLiberalize is positive but statistically insignificant, suggesting that the highest income quintile in control nations is approximately as well off as the same class in liberalizing nations. It is possible to reconcile this result with [Fig. 1a and b](#). However, in Column (2) the coefficient on NonLiberalize is positive and statistically significant, indicating that the middle class in control nations actually fares better than its counterpart in liberalizing nations. A similar result is found for the lowest income quintile group in Column (3). In both Columns (1) and (2), PostLiberalize_{*t*} is robust to the addition of the NonLiberalize indicator in both sign and magnitude.

In Panel (B) we extend the estimation idea of Section 3.2, by testing for differential slope effects in non-liberalizing versus liberalizing nations. Column (5) shows that the inclusion of the interactive variables reduces the magnitude as well as the statistical significance of NonLiberalize in the Q_m regression. Its statistical significance is retained in Column (6),

¹² Although [Fig. 1b](#) indicated no particular pattern in the change in income shares for non-liberalizing nations, that illustration was given for an arbitrary break point year (1989). Further, [Fig. 1b](#) is a useful summary of the data but does not control for country specific factors or year effects.

Table 5
Income shares in non-liberalizing nations

	(A)			(B)		
	Q_h	Q_m	Q_l	Q_h	Q_m	Q_l
	(1)	(2)	(3)	(4)	(5)	(6)
PostLiberalize	0.034* (0.019)	− 0.030* (0.016)	− 0.008 (0.006)	0.031* (0.019)	− 0.028* (0.016)	− 0.005 (0.005)
NonLiberalize	0.301 (0.193)	0.247** (0.125)	0.123* (0.069)	0.260 (0.201)	0.223* (0.131)	0.112** (0.044)
SMV_t	0.335** (0.111)	− 0.277** (0.085)	0.066 (0.052)	0.206* (0.121)	− 0.199* (0.126)	− 0.007 (0.036)
SMV_{t-1}	0.667** (0.243)	− 0.482** (0.186)	− 0.111 (0.100)	− 0.056 (0.404)	− 0.030* (0.235)	0.033 (0.055)
NonLiberalize × SMV_t				− 0.284 (0.188)	0.203* (0.089)	− 0.019 (0.054)
NonLiberalize × SMV_{t-1}				0.298 (0.732)	− 0.129 (0.123)	0.386 (0.216)
Year Effects	✓	✓	✓	✓	✓	✓
Country Dummies	✓	✓	✓	✓	✓	✓

The dependent variable is the income share of the Q_j^{th} income group ($j=h,m,l$). The explanatory variables include *NonLiberalize* which is an indicator of nations that experienced no liberalization policies, *NonLiberalize* × SMV_t which is the interaction of *NonLiberalize* and SMV_t , and *NonLiberalize* × SMV_{t-1} that is similarly defined as an interaction of its constituent variables. Country-specific dummies, year dummies, and all the control variables from Table 3a were estimated but are not reported for brevity. The reported standard errors are heteroskedasticity-consistent (White) and reported in parentheses. * and ** denote significance at the 10 and 5 percent error level, respectively.

where its magnitude is also slightly increased. Our primary interest is in the joint effect of the contemporaneous and lagged values of SMV and their interaction with NonLiberalize.

In Column (4), the results are somewhat robust to the inclusion of the interaction variable. The coefficient on SMV_t is significantly estimated at 0.206. While the coefficient on NonLiberalize × SMV_t is insignificant at conventional levels, something can still be said. The point-estimates indicate that a 10% increase in equity market activity (i) raises the upper quintile's income share in liberalizing nations¹³ by 2.06 percentage points, but (ii) decreases Q_h by 0.78 percentage points in non-liberalizing countries; a result we qualify by stating that the obtained estimates are significant only at the 85th percentile.

In Column (5), NonLiberalize × SMV_t is estimated to be positive and statistically significant, while the coefficient on the contemporaneous variable SMV_t is estimated as negative and statistically significant. These results jointly indicate that (i) a 10% increase in equity market valuation results in lowering middle class income shares in liberalizing nations by 1.9 percentage points, but (ii) the same 10% increase in valuation is associated with an increase of 0.0045 (=0.2035−0.1991) percentage points in the middle class income share in non-liberalized nations. Although this is a statistically significant result, the implied numbers are economically less significant, i.e., they suggest that the change in middle class income shares is close to zero in non-liberalizing nations. The coefficients on

¹³ The slope effect of a change in stock market valuation is given as $\delta_3 + \delta_5 SMV$, or, a δ_3 change in liberalizing nations, and a $\delta_3 + \delta_5$ effect in non-liberalizers.

the lagged value of equity valuation, SMV_{t-1} as well as its interaction with the non-liberalizing indicator are estimated imprecisely, and cannot be distinguished from zero.

Overall, we find support for the hypothesis that increases in equity market activity are associated with income shares differentially in liberalizing versus non-liberalizing nations. However, this relationship is estimated less precisely and attenuates to zero for the highest income share Q_h .¹⁴

3.4. The staggered effect of liberalization

The next set of results is reported for models styled in the standard “event-study” specification pioneered by [Eckbo \(1983\)](#). Event study models are somewhat atypical for the current context because in contrast to changes in stock price valuations, we expect financial market liberalization to have continuous, diffused effects on income and income distributions, rather than discrete jump effects along the time line. Nevertheless, in keeping with some of the prominent themes in empirical finance, we begin with a simple specification, estimating variants of the model:

$$Q_{jit} = \alpha_i + \sum_{k=0}^K \text{PostLiberalize}_{it-k} \delta_k + X_{it}' \beta + \varepsilon_{it}$$

where the variable $\text{PostLiberalize}_{it+k}$ is now redefined as follows

$$\text{PostLiberalize}_{it+k} = \begin{cases} 1 & \text{if it is year } t+k \text{ after } i\text{'s first capital account liberalization policy} \\ 0 & \text{otherwise,} \end{cases}$$

and X constitutes a set of control variables used in [Tables 3a–5](#) such as GovtConsumption, SecondaryEnrollment and Rule of Law. We replace the year dummies with a measure of regional GDP growth to account for some of the regional business cycle effects. The latter is a useful substitution, as the number of coefficients to estimate in this model increase linearly with the parameter K .¹⁵

The major statistical difference afforded by the event study analysis relative to the linear regression model is its accommodation of a staggered effect of liberalization on income shares for K periods beyond the first year of liberalization. Pooling all post-liberalization years together, as in [Bekaert et al. \(2000\)](#), is informative about mean differences before and after liberalization, but cannot separate the finer differences across these years. Although we experimented with a grid of values for $K=0, 1, 2, \dots, 6$, the reported results are for $K=4$ because the addition of every additional year beyond $K=4$ was found to have no power in explaining the variance in Q_j ($j=l, m, h$).

¹⁴ One might also directly test whether [Fig. 1b](#) is confirmed in the data by a simple regression of income shares on an indicator that takes on the value 1 for the breakpoint year or thereafter, and 0 otherwise. We find that such regressions are sensitive to the choice of the breakpoint year. Further, the breakpoint year is arbitrary and with no economic significance for these control nations so we do not report those results.

¹⁵ One point of concern is that the number of regressors in this model will grow quite fast, limiting precision. To address this concern, we pared down the set of regressors in this model to those that we deem are the most important, from [Table 3a](#). We find that making the relevant change makes our estimates more precise.

Staggering the effects of liberalization could also be useful in speculating about the source of the changes in shares. For an extreme example, suppose the upper quintile's income were completely linked with the equity market. Then, based on the evidence in Henry (2000a) that liberalization leads a rise in emerging market equity prices, one would expect instantaneous changes in their income shares in the event of liberalization, and a coefficient on PostLiberalize_t that is significantly different from zero. Alternatively, if the changes in income depend on realized profits from private investments, it is plausible that the measured response of liberalization is diffused over time, such that the coefficients on $\text{PostLiberalize}_{t+k}$ ($k=1, 2, \dots$) are different from zero, while the coefficient on PostLiberalize_t is not.

Table 6 presents the results of estimating the model above. The first row, which presents robust regression estimates of the model above, indicates that in the year of liberalization the top quintile's share (Q_h) is positively affected, being 2.2% above its non-liberalization mean. This strand of evidence is consistent with the notion that gains to the upper quintile may be realized imminently after liberalization because of the instantaneous reaction of equity markets, whose valuation directly affects holders of equity. Because increases in one share must be offset by decreases in one or more of the other income shares, we expect coefficients on Q_m and/or Q_l to be negative, which is found to be the case in Columns (2) and (3). That these coefficients are statistically indistinguishable from zero could arise for two reasons. First, sharp declines within Q_m and Q_l in some countries may be offset by huge gains in Q_m and Q_l in others, leading to no observed change in an average (i.e., regression) sense in these income shares. This explanation would be consistent with the observed increase in Q_h and the absence of an estimated decline in Q_m and Q_l . An alternate explanation is that the offsetting declines are felt in narrower quantiles of the income distribution than we consider. To explore this possibility, the regressions are re-

Table 6
Staggered effects on income shares

	(A)			(B)		
	Q_h	Q_m	Q_l	Q_{m2}	Q_{m3}	Q_{m4}
PostLiberalize_t	0.022** (0.006)	-0.010 (0.012)	-0.001 (0.002)	-0.004 (0.005)	-0.007** (0.003)	-0.002 (0.004)
$\text{PostLiberalize}_{t+1}$	0.042* (0.022)	-0.029** (0.014)	0.004* (0.002)	-0.008 (0.006)	-0.014** (0.004)	-0.007 (0.005)
$\text{PostLiberalize}_{t+2}$	0.024** (0.008)	-0.028** (0.012)	0.002 (0.002)	-0.006 (0.005)	-0.011** (0.003)	-.009** (0.004)
$\text{PostLiberalize}_{t+3}$	-0.007 (0.003)	0.003 (0.017)	0.002 (0.003)	-0.001 (0.007)	-0.001 (0.004)	0.002 (0.006)
$\text{PostLiberalize}_{t+4}$	0.004 (0.003)	-0.011 (0.017)	0.005 (0.003)	-0.005 (0.007)	-0.003 (0.004)	-0.007 (0.006)

The dependent variable is the income share of the Q_j^{th} income group ($j=h,m,l$). Q_{m2} , Q_{m3} and Q_{m4} are the second, third, and fourth quintile, respectively, whose sum constitutes Q_m . The explanatory variables PostLiberalize_t , $\text{PostLiberalize}_{t+1}$, $\text{PostLiberalize}_{t+2}$, $\text{PostLiberalize}_{t+3}$ and $\text{PostLiberalize}_{t+4}$ are as defined in Section 3. Country-specific dummies were estimated but are not reported for brevity. Standard errors are heteroskedasticity-consistent (White), and are reported in parentheses. * and ** denote significance at the 10 and 5 percent error level, respectively.

estimated on each of the three components that make up Q_m , i.e., the second, third, and fourth quintiles that we abbreviate as Q_{m2} , Q_{m3} and Q_{m4} . These results are reported in Panel (B) of Table 6.

We find that in the first year of liberalization, there is a statistically significant negative effect on the income share Q_{m3} , which is the center of the “middle class” income share, offsetting the rise in the income share of the top quintile. The point estimate suggests that the income share of this group fell 1.2 percentage points, while there was no significant effect on either Q_{m2} , or Q_{m4} . This finding reconciles the apparent lack of decline in the aggregate measure Q_m . It is also plausible that the observed gains to the upper quintile are driven by gains to a much narrower group than to the entire upper quintile; however, as there is no data on say, deciles, this hypothesis cannot be verified empirically.

The results of Table 6 indicate that while some movements in income shares are observed as early as the year of liberalization, much of the changes occur over the next 2 years and diminish thereafter. The results indicate that the upper quintile’s income share is 4.2 and 2.4 percentage points larger than the sample mean in the first and second year after liberalization respectively, as seen by the coefficients on $\text{PostLiberalize}_{t+1}$ that and $\text{PostLiberalize}_{t+2}$. Contemporaneously offsetting these gains, the middle class income share falls 2.9 and 2.8 percentage points in the first and second years after liberalization respectively. In the first year after liberalization, the lowest income share, Q_l , also falls 0.4%. It is interesting to note that in the second year after liberalization where there is no discernible effect on Q_l , the gains to Q_h exactly offsets the reduction in Q_m . In each of the third and fourth years after liberalization there is no statistically discernible conditional mean difference in any of the income shares.

3.5. Contemporaneous reforms

One concern in empirically isolating the impact of capital account liberalization is that of controlling for the contemporaneous implementation of other economic or political reforms which might systematically influence the “climate” for growth. Henry (2000a) documents that, with very few exceptions, nations that undertook capital account liberalization coincidentally implemented stabilization and trade liberalization policies, accelerated privatization, and lifted restrictions on exchange rates. King and Levine (1993), Rajan and Zingales (1998), Levine and Zervos (1998) indicate that growth is strongly associated with development in the banking sector. Without controlling for coincidental domestic reforms or world business cycle effects that are themselves associated with changes in the income distribution, a regression framework would misrepresent the association of stock market liberalization with changes in income shares, if such co-movements are inadequately captured by year dummies.

To address this concern, the regression model in Section 3.1 is supplemented with additional controls for specific domestic reforms. Because we find in Table 6 that beyond the second year after liberalization the effects are insignificant both economically and statistically, we consider only the 2 years subsequent to the year of implementing capital account liberalization. Data on the other major domestic reforms in the emerging markets are from Henry (2000a). With these data, (for $k=1, 2$), we add indicator variables for

stabilization policies (Stabilize_{t+k}), slackening of exchange rate restrictions (Exchange_{t+k}) and the implementation of privatization reforms (Privatize_{t+k}). These variables are defined analogously to $\text{PostLiberalize}_{t+k}$,¹⁶ taking on the value of 1 if it is the $(t+k)$ th year after the reform was implemented, and 0 everywhere else.

Results from estimating the modified regression model are given in Table 7. We find that relative to Table 6, the coefficient on the $\text{PostLiberalize}_{it}$ indicator becomes significant in the Q_m regression and its statistical significance is retained in the Q_h regression. Note that both these coefficients are more positive relative to their estimates in Table 6. This result is evidence that implementation of the other major domestic reforms is positively correlated to a liberalizing reform. We find that stabilization policies appear to have their strongest effect a year after they are implemented, as indicated by the significant coefficient on Stabilize_t . The results indicate that stabilization policies are in general beneficial to the lower income classes. Likewise, the effects of privatization reforms are felt with a year's lag as seen in the coefficients of Privatize_{t+1} in both Columns (1) and (2). Privatization reforms, whose effects are an order of magnitude smaller than that of stabilization programs, are positively associated with the upper quintile, and negatively with the middle class. As with both the above policies, the easing of exchange restrictions also affects income shares with a lag; the results discern strong effects in both the first and second year after more lax exchange rate restrictions are implemented as seen in the coefficients of Exchange_{t+1} and Exchange_{t+2} .

Table 7 indicates that it is important to control for contemporaneous domestic programs of stabilization, privatization policies and the easing of exchange rate restrictions. Importantly, the coefficients on the PostLiberalize indicators are statistically significant even after controlling for these other reforms, which themselves impact income shares. We infer that the implementation of the other major domestic reforms is positively correlated with capital account liberalization because of the direction in which the coefficients of the PostLiberalize indicators change relative to Table 6. In general, capital account liberalization appears to impact income shares contemporaneously and up to 2 years after liberalization, while other reforms have no contemporaneous effect.

4. Measures of stochastic and Lorenz dominance

The regression analysis presents evidence that higher income groups benefited unambiguously post-liberalization, with growths in their income share surpassing the

¹⁶ These data are from Henry (2000a). “Stabilization programs” refers to programs enacted to require stricter and more transparent monetary policy, lifting of price and exchange rate controls, steps taken to prevent federal and state banks from money printing and “large” public debt rescheduling. Analogously, “privatization” is a term reserved explicitly for actual privatization of firms, and/or steps taken to accelerate privatization practices already in place including the enactment of official laws to permit private individuals into industries earlier reserved for the government, and the transfer of ownership to private individuals, including foreign ownership. “Exchange” refers to a significant easing of foreign exchange restrictions (e.g., easier profit remittance for foreign firms in Venezuela in 1989). Additional details are documented by Henry (2000a).

Table 7

Income shares, liberalization and concurrent reforms

	Q_h	Q_m	Q_l
PostLiberalize _{<i>t</i>}	0.016** (0.008)	− 0.017** (0.006)	− 0.001 (0.002)
PostLiberalize _{<i>t</i>+1}	0.022 (0.019)	− 0.031** (0.013)	0.009 (0.006)
PostLiberalize _{<i>t</i>+2}	0.051** (0.017)	− 0.05** (0.013)	0.003 (0.006)
Stabilize _{<i>t</i>}	0.022 (0.014)	− 0.026** (0.026)	0.002 (0.005)
Stabilize _{<i>t</i>+1}	− 0.084** (0.031)	0.056* (0.027)	0.024** (0.010)
Stabilize _{<i>t</i>+2}	0.002 (0.010)	0.001 (0.010)	− 0.0079 (0.003)
Privatize _{<i>t</i>}	− 0.037 (0.020)	0.040** (0.017)	− 0.0027 (0.007)
Privatize _{<i>t</i>+1}	0.041** (0.010)	− 0.033** (0.006)	− 0.008 (0.002)
Privatize _{<i>t</i>+2}	0.030* (0.018)*	− 0.011 (0.015)	− 0.017 (0.007)
Exchange _{<i>t</i>}	− 0.015 (0.016)	0.021 (0.011)	− 0.006 (0.004)
Exchange _{<i>t</i>+1}	0.046** (0.022)	− 0.036* (0.016)	− 0.011 (0.007)
Exchange _{<i>t</i>+2}	− 0.026** (0.011)	0.022* (0.010)	0.006 (0.004)
Country Dummies	✓	✓	✓

The dependent variable is the income share of the Q_j^{th} group ($j=h,m,l$). The explanatory variables *Stabilize*, *Privatize* and *Exchange* represent indicator variables for the implementation of stabilization programs, privatization policies and weakening of exchange rate restrictions, respectively. *Stabilize_t*, *Stabilize_{t+1}*, *Stabilize_{t+2}* are variables which take a value 1 during the year of liberalization, the year after liberalization, and 2 years after liberalization respectively and 0 everywhere else. The *Privatize* and *Exchange* variables are defined analogously. Country-specific dummies, and controls for *Regional GDP growth* (from year t to $t+1$), *Govt Consumption*, *Secondary Enrollment* and *Rule of Law* are also estimated but are not reported for brevity. ** and * denote significance at the 5 and 10 percent error level, respectively, and standard errors are White-consistent.

growth in mean income, while the middle class unambiguously suffered lowered income shares. However, as discussed earlier, this result must be qualified with the fact that with very few exceptions, income *levels* grew almost uniformly in all income classes in our sample. In this section, we revert to studying income levels using standard measures of stochastic and Lorenz dominance to empirically verify the shifts in income distributions from the pre-liberalization to post-liberalization era.

The three dominance measures discussed below are first order stochastic dominance (FOSD), second-order stochastic dominance (SOSD) and Lorenz dominance, also called mean-normalized second order stochastic dominance. Since the purpose is to compare inequality before and after financial liberalization, we are primarily interested in the concept of Lorenz dominance. However, comparisons using first- and second-order stochastic dominance are also presented since these give a broad picture of the mean changes in the income distribution during the process of financial liberalization. Based on the empirical finding from the regressions, we expect (i) the post-liberalization distribution to be characterized by FOSD and SOSD relative to the pre-liberalization distribution, while (ii) Lorenz dominance should not hold for the pooled sample of liberalizing countries. We use a method known as the “p-approach to dominance”, where the measures of dominance can be expressed in terms of quantiles (Davidson and Duclos, 2000). Although income quintiles are a rough approximation to the actual income distribution, they are nevertheless useful in comparing the broad patterns of income distribution before and after liberalization.

Denote the cumulative distribution function of the inflation-adjusted incomes of the representative individuals prior to and subsequent to liberalization as F_A and F_B , respectively. F_B first order stochastically dominates (FOSD) distribution F_A if and only if $F_B(y) \leq F_A(y)$. Let I denote income. For the p th quintile, FOSD is equivalent to verifying whether $I_A(p) \leq I_B(p)$, i.e., whether the representative individual in each quintile is better off in the post-liberalization distribution than in the pre-liberalization distribution. SOSD is implied by FOSD. The test for SOSD entails verifying whether $CI_A(p) \leq CI_B(p)$, for each p , where $CI_A(p)$ and $CI_B(p)$ denote the mean cumulative income up to the p^{th} quintile of distributions A and B, respectively. Fig. 2a and b illustrate the shifts in income distribution using these two measures of dominance for the combined sample of liberalizing countries before and after financial liberalization. The evidence seems to be overwhelmingly in favor of first-order (and hence second order) stochastic dominance. A similar exercise for individual countries yields similar results with FOSD holding for 10 out of the 11 countries, and mean inflation-adjusted incomes of each of the quintiles increasing by more than 5% in the 5-year window after liberalization.

A Lorenz curve plots the cumulative income share of the bottom p -quantiles of the population against p . Distribution B Lorenz-dominates distribution A if the Lorenz curve associated with distribution B lies nowhere below and at least somewhere above that of A. Equivalently, distribution B dominates distribution A if $CF_A(p) \leq CF_B(p)$, where $CF_A(p)$ and $CF_B(p)$ are the cumulative income shares up to the p^{th} quintile of distributions A and B, respectively. Lorenz dominance can be interpreted as a shift in income distribution where the representative individual in the poorest p -quintiles receives a greater share of the income in the new state of nature after liberalization. We obtain Lorenz dominance for two and crossings for the rest nine liberalizing countries. Fig. 2c illustrates the absence of Lorenz dominance for the combined sample of liberalizing countries in the 5-year window after financial liberalization.

The simultaneous empirical observance of FOSD and the absence of Lorenz dominance reinforce our earlier conclusion that even though mean incomes increased for all the income groups post-liberalization, the gains from liberalization were not equitably distributed. The aggregate dominance results, combined with our regression analysis, make it highly likely that similar effects would be discernible at the micro-level as well for most of the liberalizing countries in our sample.

5. Can a case be made for causality?

Evidence from the regression analysis indicates that there are real statistical associations between the occurrence of liberalization and the income shares of the highest quintile and the “middle class”. These relations persist statistically in the presence of controls for domestic fundamentals, world business cycle movements and country-specific factors. They are also robust upon controlling for contemporaneously implemented domestic stabilization and privatization reforms, and the easing up of exchange restrictions. The results indicate that the highest income quintile benefits at the expense of a large middle class. The lowest quintile’s income share remains unaffected, mirroring a recent finding in Dollar and Kraay (2001).

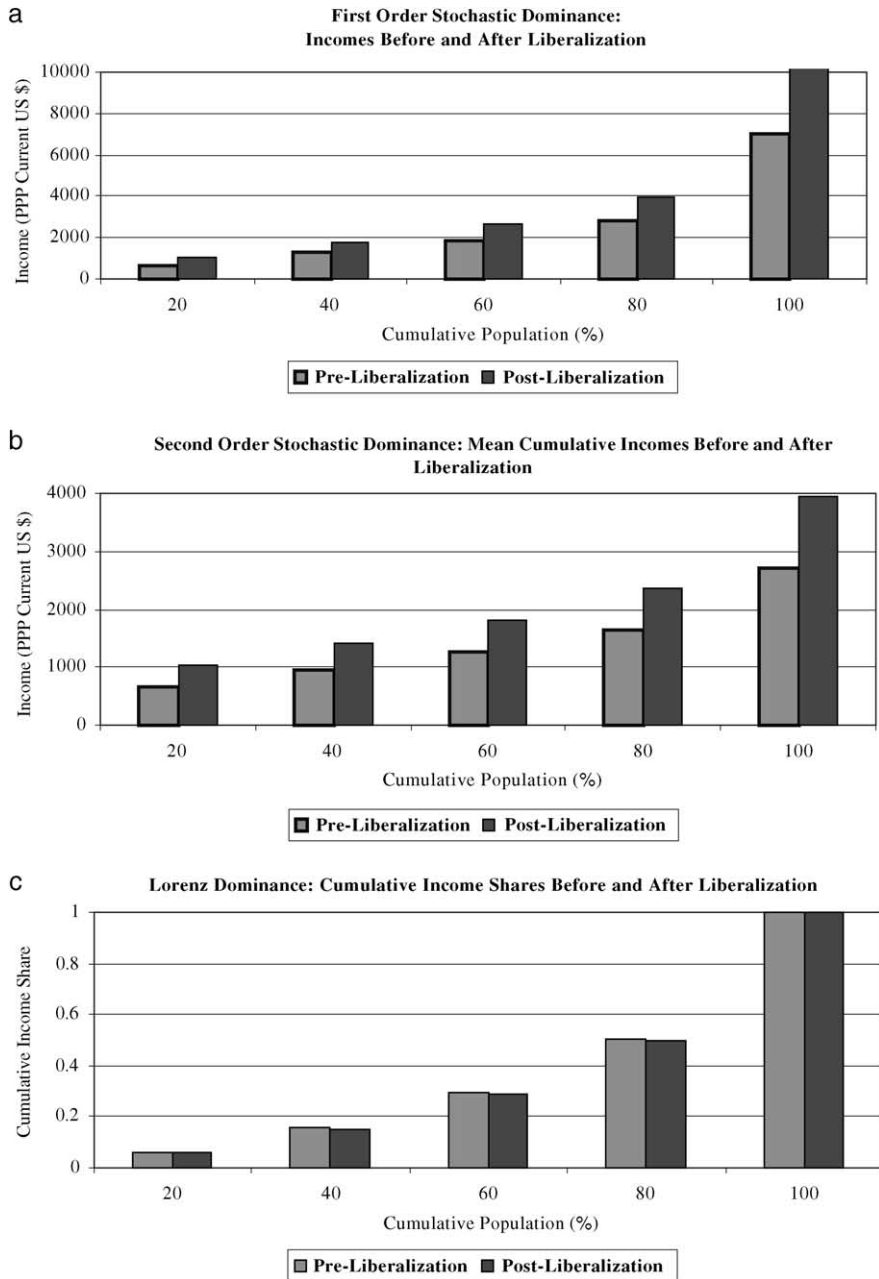


Fig. 2. (a) FOSD. (b) SOSD. (c) Lorenz Dominance.

It is worth questioning whether the obtained results are simply a result of endogenous policy reforms that lead to the usual identification problem in cross-country regressions. Consider the following scenario. Suppose members of the upper quintile accurately predict a future positive shock to the marginal product of capital and, recognizing that foreign capital flows will have beneficial price effects on equity, influence policymakers to implement liberalizing reforms. Then, one would expect a rise in the upper quintile's income share due to the positive shock to the productivity of capital, and liberalization will have no causal role for their higher income share, although liberalization will be positively correlated with the upper quintile's income share. This scenario is consistent with our empirical findings. In this case, the liberalization indicator simply absorbs the effects of an unobserved variable representing expectations. In the absence of a convincing way to control for such a possibility, we do not press on a causal role for liberalization.

To this end, we temper our findings in the following ways. First, if the decision to implement liberalizing reforms was in correct anticipation of higher equity valuations in the future, then the obtained results must be different in magnitude than the true association between liberalization and the income shares. In particular, the documented statistical association between liberalization and the upper/middle income shares may be larger than the true association.

Second, even in the absence of endogeneity, all our estimates are obtained conditional on the world business cycle effects in place at the time of liberalization. As documented by [Beim and Calomiris \(2000\)](#) and [Bekaert et al. \(2002\)](#), the late 1980s and early 1990s saw an unprecedented rise in capital flows to emerging markets, at a time when global interest rates were fairly low. The rise in capital flows was linked to the expansion of international finance fuelled by deregulation in many parts of the world as well as technological advances in computing and risk analysis which supported the creation of global financial institutions. Therefore, out-of-sample predictions for future liberalizers based on our results, must be tempered to account for future global conditions, such as high world real interest rates, which may not facilitate large net flows to emerging markets.

On a related note, there may be an element of a “first-mover” advantage for the nations that liberalized first. It is plausible that for the first set of liberalizing nations, foreign investors discounted the importance of political uncertainties, judicial systems and the fragility of certain emerging market nations¹⁷ when investing in these countries. Over time, there is likely to be more learning for international investors about the nature of risk, as well as the political and legal institutions, in emerging markets. It is conceivable that such learning result in a more cautious approach to emerging market investments in the future. Under this scenario, one may observe a much less robust

¹⁷ Chile, Brazil and Venezuela are examples of the liberalizing markets that are relatively fragile. These nations are heavily dependent on their undiversified export sectors, which makes their terms of trade highly sensitive to fluctuations in commodity prices which can lead to the deterioration of their debt or equity values (see [Beim and Calomiris, 2000](#)).

association of liberalization with incomes, or income shares than has been documented in this paper.

6. Conclusions

This paper presents evidence of a strong statistical association between the event of liberalization and income shares. The data strongly support a positive coefficient between liberalization and the highest income quintile's share of mean income, and a negative coefficient between liberalization and the middle class income share. For our study, the middle class represents the aggregate sum of the three middle quintiles. We find no evidence of any statistical association between liberalization and the lowest income quintile. Although the middle class "suffers" in the wake of a liberalizing reform while the upper quintile gains, this statement is true for income shares. We find that income levels in liberalizing nations almost universally rise after liberalization. We provide tests of stochastic dominance of the pre- and post-liberalization income distributions to complement our regression analysis.

The patterns we describe persist in the presence of a wide variety of controls for domestic fundamentals, world business cycle movements and country-specific factors. They are also robust to the addition of controls for contemporaneous domestic reforms. We find that equity revaluations affect income shares differentially before and after liberalizations. They also affect income shares differentially across liberalizing and non-liberalizing nations.

Because liberalizing and other major domestic reforms are rarely implemented in a vacuum, we do not press on causality. It is possible that endogenous policy decisions will attenuate some of the aforementioned correlations that we document. However, it is important to note that there are mechanisms which should relate capital market liberalizations to income shares under a wide variety of hypotheses that are true in emerging markets (e.g., differential access to credit markets, limited stock market participation, and the tight links between upper quintiles and policy makers). In any event, the patterns we describe should prove useful in the debate on emerging markets phenomena, and add to the research that analyzes the aftermath of capital account liberalization in emerging markets.

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Appendix A. Data sources and variable glossary

Data source and variable name	Variable description	Detailed variable description
Deiningering and Squire and WIID 2000		
Q_5	Income share of the highest quintile	Income share accruing to the highest quintile of the population. Income shares are constructed from reliable income or expenditure data referring to the (entire) national population.
Q_{234}	Income share of the middle three quintiles	Sum of income shares accruing to the middle three quintiles of the population.
Q_{m1}	Income share of the second quintile	Income share accruing to the second quintile of the population.
Q_{m2}	Income share of the third quintile	Income share accruing to the third quintile of the population.
Q_{m3}	Income share of the fourth quintile	Income share accruing to the fourth quintile of the population.
Q_1	Income share of the lowest quintile	Income share accruing to the lowest quintile of the population.
World bank database on financial structure and economic development		
SMC	Stock Market Capitalization/GDP	Stock market capitalization to GDP equals the value of listed shares divided by GDP. Both numerator and denominator are deflated appropriately, with the numerator equaling the average of the end-of-year value for year t and year $t - 1$, both deflated by the respective end-of-year CPI, and the GDP deflated by the annual value of the CPI.
SMV	Stock Market Valuation/GDP	Stock market valuation equals the value of total shares traded on the stock exchange to GDP.
Private Credit	Private credit by deposit money banks and other financial institutions to GDP	Private credit by deposit money banks and other financial institutions to GDP equals claims on the private sector by both deposit money banks and other financial institutions divided by GDP.
WDI STARS 2000 GDP	GDP per capita, PPP (current international \$)	GDP per capita based on purchasing power parity (PPP). GDP PPP is gross domestic product converted to international dollars using purchasing power parity rates. An international dollar has the same purchasing power over GDP as the US dollar in the United States.
Govt. Consumption	General government consumption (% of GDP)	General government consumption includes all current government spending for purchases of goods and services (including wages and salaries).

Appendix A (*continued*)

Data source and variable name	Variable description	Detailed variable description
Secondary Enrollment	School enrollment, secondary (% gross)	Gross enrollment ratio is the ratio of total enrollment, regardless of age, to the population of the age group that officially corresponds to the level of education shown.
Terms of trade	Terms of trade adjustment (constant LCU)	The terms of trade effect equals capacity to import less exports of goods and services in constant prices. Data are in constant local currency.
IRIS-3		
Rule of Law	Rule of law indicator	Rule of law contains annual values for the Rule of Law indicator (on a scale of 0–6) for the years 1982–1997, constructed by Stephen Knack and the IRIS Center, University of Maryland, from monthly data from the International Country Risk Guide (ICRG).

Notes: (1) WIID refers to the UNU/WIDER-UNDP World Income Inequality Database 2000. (2) WDI STARS refers to World Bank Socio-economic and Time Series Retrieval System CD-ROM 2000.

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