ESSAYS ON FINANCIAL MARKETS AND MACROECONOMICS

by

Alessandra Bonfiglioli

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ABSTRACT

This thesis consists of three papers addressing different aspects of financial markets and institutions.

Equities and Inequality studies the relationship between investor protection, the development of financial markets and income inequality. In the presence of market frictions, investor protection promotes financial development by raising confidence and reducing the costs of external financing. Developed financial systems spread risks among financiers and firms, allocating them to the agents bearing them the best. Therefore, financial development plays the twofold role of encouraging agents to undertake risky enterprises and providing them with insurance. By increasing the number of risky projects, it raises income inequality; by extending insurance to more agents, it reduces it. As a result, the relationship between financial development and income inequality is hump-shaped. Empirical evidence from a cross-section of sixty-nine countries, as well as a panel of fifty-two countries over the period 1976-2000, supports the predictions of the model.

How Does Financial Liberalization Affect Economic Growth? assesses the effects of international financial liberalization and banking crises on investments and productivity in a sample of 93 countries (at its largest) observed between 1975 and 1999. I provide empirical evidence that financial liberalization spurs productivity growth and marginally affects capital accumulation. Banking crises depress both investments and TFP. Both levels and growth rates of productivity respond to financial liberalization and banking crises. The paper also presents evidence of conditional convergence in productivity across countries. However, the speed of convergence is unaffected by financial liberalization. These results are robust to a number of econometric specifications.

Explaining Co-movements Between Stock Markets: US and Germany explains co-movements between stock markets by explicitly considering the distinction between interdependence and contagion. It proposes and implements a full information approach on data for US and Germany to provide answers to the following questions: (i) is there long-term interdependence between US and German stock markets? (ii) Is there short-term interdependence and contagion between US and German stock markets, i.e. do short-term fluctuations of the US share prices spill over to German

share prices and is such co-movement unstable over high volatility episodes? The answers are no to the former and yes to the latter.

Ai Miei Genitori

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Chapter 1

Introduction

The allocation of resources and risks in the economy, as well as the returns from assets, are to a great extent determined through the financial markets. The functioning and cross-country integration of financial systems may influence saving and investment decisions, technological innovation and occupational choices, thereby affecting the wealth of nations, its distribution and growth rate. My thesis is part of a project aimed at studying the links between institutional features of financial systems, their structure and a series of macroeconomic variables. Each of the following chapters analyzes a specific aspect of this large picture.

Chapter 2 investigates the link between investor protection, financial development and income inequality. The contribution in this chapter is both theoretical and empirical, and mainly related to the literature on financial development, growth and income distribution (see Levine, 2005 for a survey) and the recent works on law and finance (see La Porta et al., 1998 for instance). In the model, agents are risk-averse and heterogeneous in their entrepreneurial ability. They face a choice between a safe and a risky technology, and entrepreneurial ability affects the probability of success in risky project. Financial markets are subject to imperfections arising from the non-observability of output to financiers, but measures of investor protection can be adopted to amend these frictions. By promoting transparency, investor protection makes misreporting output costly for entrepreneurs. Better guarantees generate more confidence among investors, thereby making them more willing to bear risk and insure the entrepreneurs. In turn, investors can spread the individual risk by holding diversified portfolios of risky activities. As a result, financial systems with stronger investor protection allow higher degrees of risk sharing. In this context, better investor protection promotes financial development and affects income inequality in three ways. (i) It improves risk sharing, thereby reducing income volatility for a given size of the risky sector; (ii) it raises the share of the population exposed to earning risk; and (iii) it increases the reward to ability. (i) tends to reduce inequality, while it is increased by (ii) and (iii). The tension between these effects gives rise to the main result in the chapter, that income inequality is a hump-shaped function of investor protection and financial development. Any improvement upon a low-level investor protection increases risk taking more than risk sharing, thereby driving up inequality. However, when investor protection is sufficiently high, any further improvement is more effective on risk sharing than risk taking, and hence reduces income inequality. As opposed to most existing work (see Greenwood and Jovanovic, 1990), here income inequality arises even in the absence of wealth heterogeneity, due to idiosyncratic factors like ability, financial market conditions and income risk. I provide empirical evidence from a cross-section of sixty-nine countries and a panel of fifty-two countries over the period 1976-2000 in support of the theoretical results.

In **Chapter 3** I turn the attention to financial globalization and its effects on economic growth. The removal of restrictions on international capital transactions has, on some occasions, been welcome as a growth opportunity (see Bekaert et al., 2003) and on others blamed for triggering financial instability and banking crises. Yet, the ongoing debate has not addressed the impact of financial liberalization on the sources of growth. Does it affect investments in physical capital or total factor productivity (TFP), or both? If so, in which ways? This chapter is a first attempt at answering these questions. Moreover, it helps understand whether financial globalization has growth or level effects and whether it brings convergence or divergence in growth rates across countries. In particular, I separately address the effects of international financial liberalization on capital accumulation and TFP levels and growth rates. Financial liberalization may affect productivity both directly and indirectly. As a direct effect, it is expected to generate international competition for funds, thereby driving capital towards the most productive projects. Indirectly, it may foster financial development, which in turn affects growth (see Levine, 2005), but may also bring about financial instability if liberalization increases the likelihood of crises (see Aizenmann, 2002). To account for both indirect effects, I control every regression for a measure of financial development and an indicator of banking crises. I follow three methodologies to assess the effects of financial liberalization and banking crises on investments and productivity, and a fourth to address the link between liberalization and crises. I use two panel datasets of at most ninety-three countries, and a cross-section of eighty-five countries over the period 1975-1999. The main results are the following. (1) The effect of financial liberalization on TFP is positive and large in magnitude, while it is weak and non-robust on investments. (2) The impact on TFP is both on levels and growth rates, implying that financial liberalization is able to spur GDP growth in the short as well as in the long run. (3) Financial liberalization only raises the probability of minor banking crises in developed countries. (4) Banking crises harm both capital accumulation and productivity. (5) Institutional and economic development amplify the positive effects of financial liberalization on productivity and limit the damages from banking crises. (6) Neither financial liberalization nor banking crises affect the speed of convergence in TFP growth rates.

Academic economists and practitioner are interested in the effects of financial globalization not only on growth and macroeconomic performance, but also on asset prices and their co-movements across world markets. This issue is addressed in **Chapter 4**, which is co-authored with Carlo A. Favero. Measuring co-movements between stock markets is a widely debated issue, due to its implications for international portfolio diversification. The literature has shown the correlations between international equity markets to vary strongly over time. This variation may be consistent with both concepts of contagion and interdependence. While interdependence accounts for the existence of cross-market linkages, contagion consists of modifications of such linkages during turbulent periods. Identifying contagion from interdependence has important implications for the understanding of potential benefits from international portfolio diversification (see, for instance, Rigobon and Forbes, 2002). This chapter proposes a methodology to disentangle interdependence from contagion in the co-movements between stock markets and applies it to the German and US stock markets. The test for the hypothesis of "no contagion, only interdependence" consists of the full information estimation of a small co-integrated structural model, built with the LSE econometric approach (see Hendry, 1995). First, the long-run equilibria are identified by testing different possible specifications. In this case, there is only one cointegrating relationship that links the (log of) US earning-price ratio to long-term interest rates, and no evidence of long-run interdependence between the two markets. Then, the Vector Error Correction Model is used as a baseline reduced form to construct a structural model for the short-run dynamics, which allows us to assess the relative importance of interdependence and contagion. The structural model shows the effect of fluctuations of US stock market on the German stock market to be captured by a non-linear specification. Normal fluctuations in the US stock market have virtually no effect on the German market, while such an effect becomes sizeable and significant for abnormal fluctuations. Such non-linearity is clearly consistent with the relevance of contagion, since it amounts to a modification of short-run interdependence in periods of turmoil. The evidence of no long-term interdependence between US and German stock markets suggests that diversification in asset allocation may be beneficial over a long-term horizon. On the other hand, the relevance of contagion in the short-term tells that any short-term asset allocation should take into account regime-switches in the relation between international stock returns.

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Chapter 2

Equities and Inequality*

1 Introduction

A recent literature on law and finance has shown that investor protection plays a significant role in promoting the development of financial markets (see Acemoglu and Johnson, 2003, La Porta et al., 1997 and 2003, and Rajan and Zingales, 2002, among others). In particular, measures aimed at improving transparency and enforcement of financial contracts reduce the costs of outside-finance (see, for instance, Shleifer and Wolfenzon, 2002) and shift risks onto the parties that can best bear them (see Castro et al., 2004). Several works have recognized the importance of financial development for various macroeconomic variables such as growth and productivity (see, Demirgüç-Kunt and Levine, 2001 for a survey). However, this growing literature has not recognized that the changes in risk-taking behavior of investors and firms, associated with better shareholder protection, may also affect income inequality. The data suggest indeed that these variables are correlated. As shown by Table 1, for a sample of sixty-eight countries observed between 1980 and 2000, the Gini coefficient of the net income distribution is on average 10% higher (at the 5% significance level) in countries where financial markets are more developed.[1] Controlling for average

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[1] I refer to the ratio of stock market capitalization over credit to the private sector as an indicator of financial development. This ratio measures the weight of equity-finance on overall borrowings, and is well suited to capture the risk sharing function of financial development. It is frequently

Table 1
Inequality, financial development and institutions - mean comparisons

	Low Smcap	*High Smcap*	*Diff*
Gini	37.48 (1.42)	41.30 (1.85)	3.819** (2.3)
$Gini_{HC}$	44.07 (1.19)	50.10 (1.46)	6.024*** (1.88)
$investor\_pr^{(a)}$	3.79 (.46)	5.95 (.67)	2.154*** (.788)
Observations	41	27	

Note. A country is labeled *High Smcap* if its ratio of stock market capitalization over credit to the private sector is above cross-sectional average. The resutls are robust to the adoption of the median as a threshold. *Gini* coefficients refer to the distribution of net per capita income, $Gini_{HC}$ are controlled for human capital. Means and differences are reported for each variable, with standard errors in parenthesis. *** and ** indicate that the difference is positive at the 1 and 5 per cent significance level. $^{(a)}$ the sample is reduced to 18 and 24 countries with *Low* and *High Smcap*, respectively. Sample period is 1980-2000.

human capital, one of the most important determinants of inequality, this difference rises to 14% (now significant at the 1% significance level).[2] Table 1 also shows that countries with more developed financial markets tend to have better institutions aimed at investor protection.[3]

This paper investigates the link between investor protection, financial development and inequality, both theoretically and empirically. It proposes a simple model where investor protection promotes financial development, thereby improving risk sharing. This induces more risk-taking in the economy and better insurance on individual earnings, which affect income inequality in opposite ways. The relationships predicted by the model are confronted with the data.

To formalize these ideas, I construct a general equilibrium two-period overlapping generations model. Agents are risk averse and heterogeneous in their entrepreneurial ability. They face a choice between a safe and a risky technology, and entrepreneurial ability affects the probability of success in risky project. I assume that financial markets are subject to imperfections arising from the non-observability

used for this purpose in the literature (see Rajan and Zingales, 2002).

[2] $Gini_{HC}$ in Table 1 is $Gini- \hat{\beta}HC$, where $\hat{\beta}$ is the OLS estimate from the regression: $Gini_i = \alpha + \beta HC_i + \epsilon_i$. HC is human capital, proxied by the share of the population aged above 25 with some secondary education (from Barro and Lee, 2001). The results do not change if I also control for the Kuznets' hypothesis by including real per capita GDP and its square, and for geography by including dummy variables. These results are available upon request.

[3] The index of investor protection is taken from La Porta et al. (2003) and accounts for measures aimed at transparency (accounting and disclosure requirements) and the enforcement of private contracts.

of output to financiers and that measures of investor protection can be adopted to amend these frictions. In particular, by promoting transparency, investor protection makes misreporting output costly for entrepreneurs.[4] For instance, this cost can be thought of as the extra-compensation the advisory firm charges to certify a falsified book. Better guarantees generate more confidence among investors, thereby making them more willing to bear risk and insure the entrepreneurs. In turn, investors can spread the individual risk by holding diversified portfolios of risky activities. As a result, financial systems with stronger investor protection allow higher degrees of risk sharing. Finally, I rule out wealth heterogeneity, so that all inequality is due to idiosyncratic factors (ability), financial market conditions and income risk. Under these assumptions, better investor protection promotes financial development and affects income inequality in three ways. (i) It improves risk sharing, thereby reducing income volatility for a given size of the risky sector; (ii) it raises the share of the population exposed to earning risk; and (iii) it increases the reward to ability. (i) tends to reduce inequality, while it is increased by (ii) and (iii).

The main result of the paper is that income inequality is a hump-shaped function of investor protection and financial development. Any improvement upon a low level investor protection increases risk taking more than risk sharing, thereby driving inequality up. However, when investor protection is sufficiently high, any further improvement is more effective on risk sharing than risk taking, hence reduces income inequality.

To make the predictions of the model more easily testable, I assume that there are only two financial instruments, which I label equity and debt. Equity makes risk sharing between investors and entrepreneurs possible, depending on the degree of investor protection, while debt does not.[5] In this way, financial development is captured by the thickness of the equity market, which is also a common empirical measure of financial development (see Rajan and Zingales, 2002, among others). Then, the testable predictions of the model will be that (1) stock market size grows with investor protection, (2) there is a hump-shaped relationship between income inequality and the thickness of the equity market, and (3) investor protection affects

[4]Also in Aghion et al. (2005), Castro et al. (2004) and Lacker and Weinberg (1989) does investor protection take the form of a hiding cost. In this paper, like in the two latter, the cost is proportional to the hidden amount, while in the first, it equals a fraction of the initial investment.

[5]This labeling is based on the common distinction between standard equity and debt contracts. However, as the financial structure becomes more developed, a variety of sophisticated debt contracts are offered to also achieve better risk sharing. These instruments, like venture capital, for instance, can be assimilated to equity in the model.

inequality only through stock market development. I provide empirical evidence from a cross-section of sixty-nine countries and a panel of fifty-two countries over the period 1976-2000 in support of these results.

The contribution of this paper is related to three main strands of literature. Acemoglu and Johnson (2003), as well as La Porta et al. (1997, 1998, 1999, 2003), show that institutions aimed at contracting protection (such as those measured by *investor_pr* in Table 1) promote the development of stock markets, but have controversial effects on economic performance. None of these studies has considered income inequality.

Many papers (Beck and Levine, 2002, Levine, 2002, Levine and Zervos, 1998, Rajan and Zingales, 1998 among others) provide empirical evidence on the link between financial development and macroeconomic variables, such as growth, investments and productivity, but none of them has addressed distributional issues.[6]

Theoretical contributions by Aghion and Bolton (1997), Banerjee and Newman (1993), Galor and Zeira (1993), Greenwood and Jovanovic (1990), and Piketty (1997), among others, have proposed explanations for the relationship between financial development, inequality and growth. In most of these models, income inequality originates from heterogeneity in the initial wealth distribution, paired with credit market frictions. As the poorest are subject to credit constraints, they are prevented from making efficient investments in the most productive activities.[7] Over time, capital accumulation determines the dynamics of wealth and income. I depart from this approach by focusing on a different source of ex-ante heterogeneity, namely entrepreneurial ability, and by describing a new mechanism translating differences in ability into income inequality that is independent of accumulation. In particular, I assume productivity to be a function of ability and that entrepreneurs have no wealth for starting their firms. By encouraging investors to ensure entrepreneurs, better investor protection allows the more talented to undertake risky projects, whose payoffs depend on ability. Heterogeneity in productivity, the extent of risk sharing and the size of the risky sector ultimately determine the income distribution. In this respect, the approach of Acemoglu and Zilibotti (1999) is closer to mine. In their paper, income inequality is generated by managerial incentives,

[6] All these works account for the influence of the legal environment on financial structure. In particular, financial variables are instrumented with legal origins, which Acemoglu and Johnson (2003) and La Porta et al. (1997, 1998, 1999, 2003) used as instruments for contracting protection.

[7] The credit constraint can derive from the non-observability of physical output as in Banerjee and Newman (1992) and Galor and Zeira (1993), or effort as in Aghion and Bolton (1997) and Piketty (1997).

which depend on risk sharing, not by ex-ante wealth heterogeneity. There, risk sharing evolves endogenously over time as a consequence of information accumulation, while here it varies only as an effect of exogenous changes in investor protection.

The only empirical assessments of the relationship between financial development and income inequality are, to my knowledge, Clarke et al. (2003) and Beck et al. (2004). Both find evidence of a negative, though non robust, relationship between the degree of financial intermediation and income inequality. The main difference with respect to my empirical analysis lays in the measure of financial development. Instead of financial depth, I use the size of the equity market relative to total credit, which seems better suited to account for the degree of risk sharing allowed by a financial system.

The remainder of the paper is organized as follows. Section 2 presents the model and its solution in partial equilibrium (a small open economy). In section 3, I study analytically and by means of numerical solution how income inequality varies with investor protection and financial development. Section 4 argues that the main results hold in general equilibrium (a closed economy). This version of the model is extensively described in the appendix. Section 5 shows that empirical evidence from a cross-section of sixty-nine countries and a panel of fifty-two countries over the period 1976-2000 supports the main results of the model. Section 6 concludes.

2 The model

2.1 Set up

The model economy is populated by two-period overlapping generations of risk-averse agents. There is no population growth and the measure of each cohort is normalized to one. For simplicity, preferences are represented by the following utility function:

$$U_t = \log(c_t) + \beta \log(c_{t+1}).$$

Second-period utility is discounted at the rate $\beta \in (0, 1)$.

At any time t, each young agent in group i is born with no wealth and ability $\pi_i \in [0, 1]$, drawn from distribution $G(\pi)$. Each group is populated by a continuum $g(\pi)$ of individuals. In the first period, agents work as self-employed entrepreneurs producing an intermediate good, and allocate their income among consumption and savings, $s(\cdot)$. When old, they invest their savings and consume all the returns before

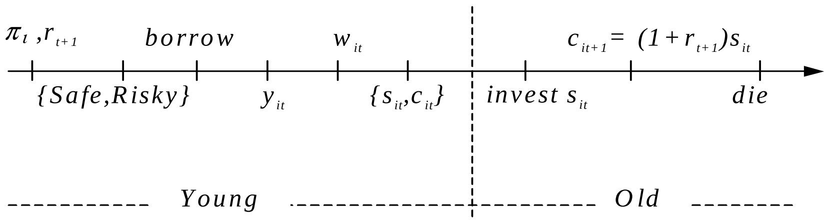

Figure 2.1: Timing of the model

dying. When investing, they can choose between safe loans, yielding a return r_{t+1}, and portfolios of risky assets. There are no bequests.

2.1.1 Intermediate goods sector

Two production processes are available to each young agent: a safe and a risky one. The safe technology does not employ capital, while the risky one requires a fixed unit investment. Therefore, the individual technological choice is analogous to an occupational choice whereby some agents become "workers" and others "entrepreneurs". In line with empirical findings, I assume that the risky activity, if successful, has higher returns than the safe one and that the probability of success depends on the ability of the entrepreneur.[8] For simplicity, and without much loss of generality, I assume that ability only affects the probability of success and not the payoffs.[9] In particular, production is given by:

$$x_{it} = \begin{cases} B & & \text{for } i \text{ running } \textit{Safe} \text{ technology} \\ A & \text{with prob. } \pi_i & \multirow{2}{*}{\text{for } i \text{ running } \textit{Risky} \text{ technology,}} \\ \varphi A & \text{with prob. } 1-\pi_i & \end{cases}$$

where $B < A$, $\varphi \in (0,1)$ and success is i.i.d. within each group. It follows that there is no aggregate risk and total production of group i equals $g(\pi_i)\,B$ or $g(\pi_i)\,[\pi_i + (1-\pi_i)\varphi A]$, depending on the technology, safe or risky, in use.

[8] See Schiller and Crewson (1997), and Fairly and Robb (2003) for empirical studies on the determinants of entrepreneurial success, mainly among small firms.

[9] Ability can be considered as playing a twofold role. It enhances the chance of succeeding in risky enterprises, as assumed in the model. But it may also raise productivity regardless of the riskiness of projects. Introducing this second effect into the model would not affect the results.

2.1.2 Final good sector

A homogeneous final good Y, used for consumption and investment, is produced by competitive firms using capital and intermediate goods. The intermediate goods produced by all agents are perfect substitutes in production. The aggregate technology has the following Cobb-Douglas form:

$$Y_t = K_{Yt}^{\alpha} X_t^{1-\alpha}, \tag{2.1}$$

where X_t is the total amount of intermediate goods, with a unit price of χ_t, and K_{Yt} is capital employed in the final good sector. Y_t is the numeraire.

2.1.3 Financial sector

Both final good firms and risky entrepreneurs need to borrow capital from the old to produce. Information about technology (A, B, φ, α) and individual ability (π_i) is public, but outside financiers cannot observe the outcome of risky activities, x_{it}. Two financial instruments, equity and debt, are available.

Equity is modeled as follows. Upon receiving one unit of capital, each young in group i commits to pay, after production, dividend payouts θ_{it}^h and θ_{it}^l in case of success and failure, respectively. Once production has occurred, unlucky entrepreneurs can only return the promised amount $\theta_{it}^l x_{it}^l \chi_t$. Successful entrepreneurs, instead, may misreport their realization of x_{it} and pay $\theta_{it}^l x_{it}^l \chi_t$, pretending to be in the bad state. However, I assume that measures of shareholder protection make misreporting costly. For every unit of hidden cash flow, the entrepreneur incurs a cost $p \in [0, 1]$. Since both ability and technology are common knowledge, either the entire $\left(x_{it}^h - x_{it}^l\right)$ or nothing is hidden, so that the payoff from misreporting is $\left(x_{it}^h - \theta_{it}^l\, x_{it}^l\right) \chi_t - p\left(x_{it}^h - x_{it}^l\right) \chi_t$. Truth-telling is rational as long as its value is at least equal to that of misreporting. Therefore, the equity contract $\left\{\theta_{it}^h, \theta_{it}^l\right\}$ must satisfy the incentive compatibility (IC) constraint:

$$v\left[\left(1-\theta_{it}^h\right) x_{it}^h \chi_t, r_{t+1}\right] \geq v\left[\left(x_{it}^h - \theta_{it}^l x_{it}^l\right) \chi_t - p\left(x_{it}^h - x_{it}^l\right) \chi_t, r_{t+1}\right], \tag{IC}$$

where $v\left[w_t, r_{t+1}\right]$ is the indirect utility of a young agent with a given income w_t and facing an interest rate r_{t+1} when old.

Debt requires a fixed repayment, R_{it}. In case entrepreneurs are not able (or willing) to pay, bankruptcy enables creditors to obtain $\min \{R_{it}\ ,\ x_{it}\}$.[10] Due to log-utility, agents in the risky sector can afford debt financing only as long as output in the bad state is higher than the interest rate. This implies that debt is always repaid and its return equals that of safe loans ($R_{it} = r_t$ for any i).

Financial contracts are set to maximize the agents' expected indirect utility, V_{it}, subject to the IC constraint and the outsiders' participation constraint. As for the latter, old agents must be indifferent between the following investments: a portfolio with shares of all group-i firms and safe loans. Risk aversion implies that debt is never optimal for financing risky projects. Furthermore, assuming that firms bear an infinitesimal cost of issuing equity, debt is preferred by the safe firms in the final good sector. Thus, payoffs from the risky technology are determined as the solution to the contracting problem for equities:

$$\max_{\theta_{it}^h,\theta_{it}^j} V_{it} \equiv \left\{\pi_i v\left[\left(1-\theta_{it}^h\right)A\chi_t, r_{t+1}\right] + (1-\pi_i)\, v\left[\left(1-\theta_{it}^l\right)\varphi A\chi_t, r_{t+1}\right]\right\}, \tag{P1}$$

subject to the incentive compatibility constraint:

$$v\left[\left(1-\theta_{it}^h\right)A\chi_t, r_{t+1}\right] \geq v\left[\left(1-\varphi\theta_{it}^l\right)A\chi_t - p\,(1-\varphi)\,A\chi_t, r_{t+1}\right], \tag{IC'}$$

and the old's participation constraint:

$$\pi_i\theta_{it}^h A\chi_t + (1-\pi_i)\,\theta_{it}^l\varphi A\chi_t = r_t. \tag{PC}$$

Note that a pooled portfolio of i.i.d. shares of group i yields the LHS of (PC) with certainty, so that the old face no uncertainty.[11]

2.1.4 Equilibrium

Firms in the final good sector are perfectly competitive and maximize profits taking prices (r_t, χ_t) as given. Each young agent from group i has perfect foresight and chooses how much to save, $s(\cdot)$, and the technology to use (safe or risky), to maximize his expected utility. Thus, each of them solves the following program:

$$\max_{T\in\{Safe,Risky\}} V_{it}^T, \tag{P2}$$

[10] Limited liability can hardly apply in this context, since the entire capital accrues to the outside financiers. Entrepreneurs do not own, and are not entitled to anything before repaying their debt.

[11] It follows that the participation constraint is the same as in the case of risk-neutral financiers with a single equity-i issuer.

where

$$
\begin{aligned}
V_{it}^{Safe} &= v\left(B\chi_t, r_{t+1}\right) \\
V_{it}^{Risky} &= \pi_i v\left[\left(1-\theta_{it}^h\right) A\chi_t, r_{t+1}\right] + \left(1-\pi_i\right) v\left[\left(1-\theta_{it}^l\right)\varphi A\chi_t, r_{t+1}\right] \\
v\left(w_{it}, r_{t+1}\right) &= \log\left[w_{it} - s\left(w_{it}, r_{t+1}\right)\right] + \beta\log\left[\left(1+r_{t+1}\right) s\left(w_{it}, r_{t+1}\right)\right] \\
s\left(w_{it}, r_{t+1}\right) &= \arg\max_{s_{it}}\left\{\log\left(w_{it} - s_{it}\right) + \beta\log\left[\left(1+r_{t+1}\right) s_{it}\right]\right\}.
\end{aligned}
$$

where w_{it} is realized income, i.e., $B\chi_t$ in case the safe technology is chosen, otherwise $\left(1-\theta_{it}^h\right) A\chi_t$ and $\left(1-\theta_{it}^l\right)\varphi A\chi_t$ in the good and bad state respectively. In other words, young entrepreneurs choose technology, given their individual ability π_i, factor prices r_t and χ_t, and the dividend payouts $\{\theta_{it}^l,\ \theta_{it}^h\}$ which solve $(P1)$.

To state the mechanism of the model in the clearest way, I first assume this to be a small open economy.[12] Both capital and intermediate goods are internationally traded, so that r_t and χ_t are exogenously given from the world markets, while Y is non traded.[13] Assuming that prices are constant, the economy is always in a steady-state and I can drop all the time indexes. For simplicity, I normalize the price of intermediate goods to one ($\chi = 1$). It follows that aggregate domestic consumption is $C = (1+\ r)K^d + \int_0^1 w\left(\pi\right) g\left(\pi\right) d\left(\pi\right)$, where K^d denotes aggregate domestic capital.

Definition *Given the interest rate r and the intermediate good price $\chi = 1$, the equilibrium for this small open economy is defined as the set of savings, technological choices and dividends $\{s_i,\ T_i,\ \theta_i^l,\ \theta_i^h\}_{i\in[0,1]}$, such that each agent in group i solves (P1) -(P2); and the factor employments $\{K_Y,\ X\}$ that maximize profits in the final good sector.*

For simplicity, I assume that $\varphi A < B+\ r < A$ and $\varphi A < r$. This implies that both safe and risky intermediate projects are run in equilibrium; and when investor protection is absent, nobody chooses the risky technology.[14]

---

[12] Later on, I will endogenize interest rate and prices, and show that the main results continue to hold.

[13] This assumption is immaterial, since factor prices are equalized everywhere.

[14] This assumption also rules out risky debt. However, it can be shown that removing this restriction would not have any considerable effect on the results.

2.2 Solution

2.2.1 Final good sector

Profit maximization by competitive firms in the final good sector yields the following demand functions for capital and intermediates: $K_Y = \alpha\frac{Y}{r}$ and $X = (1-\alpha)Y$. Market clearing requires $Y = C + K^d$.

2.2.2 Young agents

Due to log-utility, the optimal saving function of each young agent is simply a constant fraction $(1+\beta)^{-1}$ of her earnings. To solve for the optimal occupational choice $(P2)$, an agent born in group i needs to know the payoffs from the risky technology. Therefore, I proceed backwards. First, I derive the optimal equity contracts $\{\theta_i^h, \theta_i^l\}_{i\in[0,1]}$ from $(P1)$, under both perfect and imperfect investor protection. Then, I characterize the occupational choice, $\{T_i\}_{i\in[0,1]}$, given the optimal payoffs. Finally, I show how the equilibrium is affected by investor protection.

Optimal equity contract: efficient markets, $p = 1$

In this case, the payoff from hiding cash flow equals earnings in the bad state, $\left(1-\theta_i^l\right)x_i^l$. This means that there is no incentive for entrepreneurs to misreport, so that investors can act as if they had perfect information about x_i. Having a state-invariant income is the first best for risk-averse entrepreneurs. Since outside financiers behave as if they were risk-neutral and perfectly informed, they are willing to provide insiders with full insurance, given that the expected return equals the safe rate. Analytically, the first-order conditions for $(P1)$ subject to (PC) require:

$$
\begin{aligned}
v_h' &= v_l' \qquad\qquad \text{and} \\
\left(1-\theta_i^h\right) &= \left[\pi_i + (1-\pi_i)\,\varphi\right] - \frac{r}{A},
\end{aligned}
$$

where v_h' and v_l' are the derivatives of $v\left[\left(1-\theta_i^h\right)A, r\right]$ and $v\left[\left(1-\theta_i^l\right)\varphi A, r\right]$ with respect to θ_i^h and θ_i^l, respectively. This means that (IC') holds with equality and $\left(1-\theta_i^h\right)A = \left(1-\theta_i^l\right)\varphi A$ (i.e., earnings of entrepreneurs are state invariant: $w_i^h = w_i^l$).

Optimal equity contract: general case, $0 < p < 1$

If investor protection is not perfect, state invariant earnings are not incentive compatible: entrepreneurs in the good state would be tempted to misreport x_i and enjoy the higher utility given by earnings $\left(1-\varphi\theta_{it}^l\right)A -\ p\,(1-\varphi)\,A$. Investors are aware of this and hence account for it when determining the dividend payouts. In other words, both (IC') and (PC) must hold with equality, so that

$$\begin{aligned} w_i^l &= \left(1-\theta_i^l\right)\varphi A = \left\{\left[\pi_i + (1-\pi_i)\varphi\right] - \pi_i (1-p)(1-\varphi)\right\} A - r, \\ w_i^h &= \left(1-\theta_i^h\right) A = \left[\left(1-\theta_i^l\right)\varphi + (1-p)(1-\varphi)\right] A. \end{aligned}$$

The wedge between state-contingent earnings, i.e. the price for the temptation to misreport, is decreasing in investor protection. If the cost of hiding profits is high, temptation to misreport is low, as is its price in terms of distance from the first best. The ratio between payoffs and ability is lower than in the efficient case, and increasing in p. This means that, by discouraging misbehavior, investor protection also fosters meritocracy. Expected earnings for entrepreneurs are the same as under perfect investor protection, but expected utility is lower, due to risk aversion. Notice that for $p = 0$, the payoffs from equity-finance are the same as those implied by a standard debt contract.

Technological choice

The solution to $(P2)$ features a threshold ability level π^* such that the *Risky* technology is chosen by any agent with ability higher than π^*. This property is formalized in Lemma 1.

Lemma 1 *There exists a unique π^* such that $\forall \pi_i \geq \pi^*$, $\pi_i v[(1-\ \theta_i^h)A, r] +\ (1-\pi_i) v\left[(1-\ \theta_i^l)\varphi A, r\right] \geq B$, and $\left\{\theta_i^h, \theta_i^l\right\}$ is the solution to $(P1)$.*

Proof. *See the Appendix.* ■

2.2.3 Investor protection and the equilibrium

Since the dividend payouts $\left\{\theta_i^h, \theta_i^l\right\}$ are functions of investor protection, also the threshold ability π^* varies with p, as formalized in Lemma 2

Lemma 2 *The threshold ability π^* is a decreasing, convex function of investor protection p.*

Proof. *See the Appendix.* ■

Given that the risky technology is financed with equity, the measure of agents who choose it represents the size of the stock market. From Lemmas 1 and 2, it follows that stock market size is a function of investor protection, as stated by Proposition 1.

Proposition 1 *Stock market size, $sm \equiv 1 - G(\pi^*)$, is increasing in investor protection, and concave for high p.*

Proof. *See the Appendix.* ■

Corollary 1 *Stock market size as a ratio of GDP is increasing in investor protection and concave for high p.*

Proof. *See the Appendix.* ■

In the efficient case ($p = 1$), the value of producing with the risky technology is higher whenever $[\pi_i + (1 - \pi_i)\varphi] A - r \geq B$. Therefore, I can easily get a closed form solution for the threshold ability,

$$\pi^*_{p=1} = \frac{(B - A\varphi) + r}{(1 - \varphi) A},$$

and verify that it lies in the support of π under the hypotheses that $A > B + r$ and $\varphi A < B + r$.

In the general case of imperfect investor protection ($p < 1$), the expression for the threshold ability is more complicated. However, payoffs are easily derived:

$$w(\pi_i) = \begin{cases} B & \text{with probability } 1 & \text{for } \pi_i < \pi^* \\ w_i^h & \text{with probability } \pi_i & \text{for } \pi_i \geq \pi^* \\ w_i^l & \text{with probability } 1 - \pi_i & \text{for } \pi_i \geq \pi^* \end{cases}$$

$$w_i^h = [\pi_i p (1 - \varphi) + \varphi + (1 - p)(1 - \varphi)] A - r \tag{2.2}$$

$$w_i^l = [\pi_i p (1 - \varphi) + \varphi] A - r. \tag{2.3}$$

Henceforth, I denote the threshold abilities associated with $p = 1$ and $0 < p < 1$ by $\pi^*_{p=1}$ and $\pi^*_{p<1}$, respectively. For $p = 1$, perfect risk sharing is achieved through equity financing so that entrepreneurs act as if they were risk-neutral. They choose the risky technology as soon as their ability implies expected earnings equal to the safe ones, i.e. $\pi_i = \pi^*_{p=1}$. This means that their earnings are state invariant and exhibit no discontinuity at the threshold ability level. When $0 < p < 1$, at $\pi_i = \pi^*_{p<1}$ the expected productivity of the risky technology needs to be higher than the productivity of the safe technology, because entrepreneurs are risk averse and cannot be fully insured through equity.

Figure 2.2 illustrates the optimal ability-earnings profiles. If there is no investor protection, nobody chooses the risky technology and hence earnings are flat and equal to B. In the opposite extreme case of $p = 1$, income of young agents is described

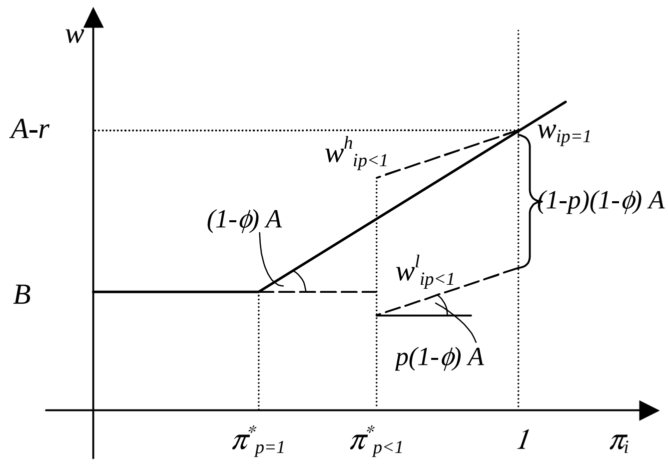

Figure 2.2: Ability-earnings profiles.

by the solid line. It is flat for the less able, who run the safe project, and proportional to ability for the more talented, risky entrepreneurs. Due to perfect risk-sharing, earnings are state invariant. If investor protection drops to $0 < p < 1$ (dashed line), equity-finance becomes more costly, thereby inducing the least able among risky entrepreneurs to shift to the safe sector. Graphically, (1) the stock market shrinks, i.e., the flat portion of the earnings profile becomes longer. I define this as the "market size" effect. (2) Proportionality between stochastic payoffs and ability becomes weaker due to higher incentives to misreport, and the wedge between state contingent earnings widens due to worse risk-sharing. I call this, as illustrated by the flatter slope and higher distance between $w^h_{ip<1}$ and $w^l_{ip<1}$, the "risk sharing" effect. The extent of imperfect risk sharing is captured by the jump in expected earnings at $\pi^*_{p<1}$. At any $\pi_i \geq \pi^*_{p<1}$, the expected payoff from the risky technology is independent of p since, for a given interest rate, the old are indifferent between stocks and loans. However, even though expected earnings are invariant, welfare is higher under perfect investor protection because of risk aversion.

3 Evaluating income inequality

In this section, I derive the key implications of the model on the overall effect of investor protection on income inequality, through the development of the stock market. To do so, I compute the variance of earnings,

$$\begin{aligned} Var\left(w\right) &= G\left(\pi^*\right)\left[B - E\left(w\right)\right]^2 + \int_{\pi^*}^{1} \left\{ \pi \left[w^h\left(\pi\right) - E\left(w\right)\right]^2 \right. \\ & \left. + \left(1-\pi\right)\left[w^l\left(\pi\right) - E\left(w\right)\right]^2 \right\} g\left(\pi\right) d\pi, \end{aligned}$$

with $E\left(w\right) = G\left(\pi^*\right) B + A \int_{\pi^*}^{1} [\pi + (1-\pi)\varphi] g\left(\pi\right) d\pi - \left[1 - G\left(\pi^*\right)\right] r$, and study how it varies with p.[15]

If there is no investor protection, all agents choose the safe technology and thus, the variance is zero. If the cost of hiding cash flow becomes any higher than zero ($p{=}\varepsilon$), some agents prefer the risky technology and raise funds through equity, thereby driving the stock market size from zero to $sm(\varepsilon)$. By the "market size" effect, a share of the economy becomes subject to income risk (having state-contingent earnings), thereby raising the variance of income (analytically, positive terms fall under the integral). Moreover, average earnings grow higher than B, so that also the agents on the flat portion in Figure 2 contribute to raising the variance.

As investor protection improves, "market size" is paired with the "risk sharing" effect, which shrinks the wedge between state-contingent earnings and hence, tends to reduce the variance. Analytically, the "risk sharing" effect tends to reduce the term under integration. The extent of the "market size" effect is decreasing in investor protection, due to the concavity of sm at high p. On the other hand, risk-sharing becomes more effective, the larger is the share of equity-financed agents. This means that, when investor protection is weak (sm is small), the market-size effect dominates because risk-sharing applies to a small fraction of the economy. Therefore, inequality at first increases with p (and with sm).

When investor protection is perfect, $Var\left(w\right) = G\left(\pi^*_{p=1}\right)\left[B - E\left(w\right)\right]^2 + \int_{\pi^*_{p=1}}^{1} \{[\pi + (1-\pi)\varphi]A - r - E\left(w\right)\}^2 g\left(\pi\right) d\pi > 0$. As p falls any lower than 1 ($p = 1-\varepsilon$), the "market size" effect drives only few agents out of the risky sector, thereby reducing income inequality by a small amount, since the difference between B, $w^h\left(\pi^*\right)$ and $w^l\left(\pi^*\right)$ is still slight. The "risk sharing" effect, instead, applies to a large share of the population, and outweighs the "market size" effect, so that there is an increase in income inequality. Therefore, improvements upon an already very good investor protection may in fact reduce inequality, although never below the case of no investor protection. Lemma 3 and Proposition 2 formalize this intuition.

[15] Since income of the old is 1-to-1+r linked to that of the young, I focus on the earnings of the active population only.

Lemma 3 *The variance of earnings is a non-monotonic function of investor protection:* $\frac{dVar(w)}{dp} > 0$ *in a neighborhood of* $p = 0$*, and* $\frac{dVar(w)}{dp} < 0$ *in a neighborhood of* $p = 1$*.*

Proof. *See the Appendix.* ■

Since, from Proposition 1, sm is continuous and monotonic in p, also the relationship between stock market size and income inequality follows a non-monotonic pattern.

Proposition 2 *The relationship between earnings variance and stock market size,* $sm \equiv 1 - G(\pi^*)$*, is non-monotonic:* $\frac{dVar(w)}{dsm} > 0$ *in a neighborhood of* $sm(0)$*, and* $\frac{dVar(w)}{dsm} < 0$ *in a neighborhood of* $sm(1)$*.*

Proof. *See the Appendix.* ■

Proposition 2 shows that income inequality, as measured by the earnings variance, increases with stock market size for small sm and falls with large sm. However, this does not give a full characterization of the relationship between inequality and stock market size for any p. Moreover, there are alternative measures of inequality, such as the Gini coefficient, that are more commonly used in empirical work. Since a characterization of this indicator is awkward to derive analytically, I obtain it through numerical solution. This exercise allows me to study the relationship between investor protection, stock market size and income inequality on the whole domain of p and to obtain a more testable version of the prediction in Proposition 2.[16]

To simulate the model, I choose parameter values consistently with the restrictions imposed on parameters throughout the paper.[17] I approximate the distribution of ability with a Lognormal(μ,σ) and parametrize the mean and variance of the associated Normal distribution, μ and σ, with values from the actual data. Although ability per se is difficult to measure, it is likely to be reflected in educational attainment. Therefore, I take the sample mean and variance of school years from the Barro and Lee (2000) database of 138 countries in 1995. Since the support of the Lognormal distribution is unbounded from above, it must be truncated to comply

[16] If the assumption that risky output in the bad state is lower than the international interest rate is removed, some of the most able agents can finance the risky project through debt, even at *p=0*. This means that the upper bound for the threshold ability becomes $\tilde{\pi} < 1$ s.t. $\tilde{\pi}v(A - r) + (1 - \tilde{\pi})v(\varphi A - r) = v(B)$, and stock market size is $G(\tilde{\pi}) - G(\pi^*)$. All results, after this relabeling.

[17] Notice that this numerical solution is for qualitative rather than quantitative purposes. Therefore, the technological parameters are not calibrated to the actual data.

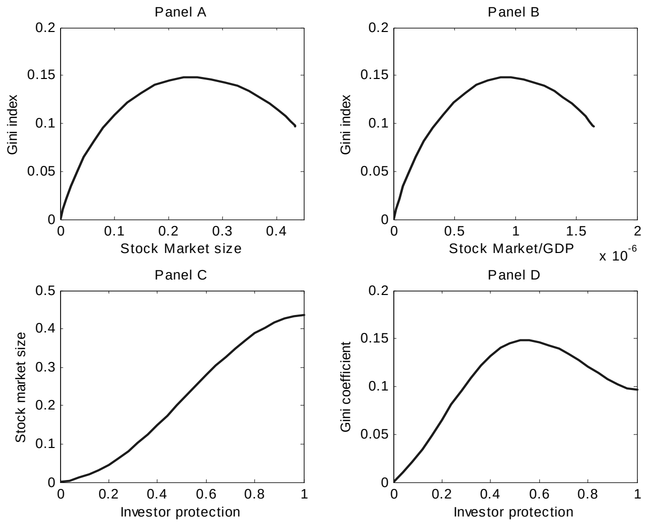

Figure 2.3: Stock market size and income inequality (Panels A-B), investor protection, stock market size (Panel C), and income inequality (Panel D). Simulation output.

with the set-up of the model. I assume the top 0.05 per cent to have ability 1, while π is lognormally distributed across the remaining 99.95 per cent of the population. I parameterize μ and σ to match the US data, where the average years of schooling are 14.258, with a variance of 26.93. I normalize the resulting ability distribution so that it fits in the interval $[0, 1]$, consistent with the model. I set $\alpha = 0.33$, $r = 0.06$, $B = 1$, $A = 2.33$, $\varphi = 0.026$, implying $sm(p = 1) \simeq 0.4$.

Both the market size and the risk-sharing effects are expected to affect the Gini coefficients and the variance of earnings in similar ways. Panel A of Figure 2.3, plotting the Gini coefficient against stock market size, confirms the expectations: the Gini exhibits a non-monotonic pattern, featuring a hump peaking at a high sm. From Corollary 1, stock market size as a ratio of GDP is monotonically increasing in investor protection, and is concave for high p. Therefore, a pattern close to Panel A can be expected for the relationship between $\frac{sm}{Y}$ and income inequality. Panel B confirms this prediction. Panel C shows stock market size to be a function of investor protection, with the properties predicted by Proposition 1. Finally, Lemma 3 is given graphical representation in Panel D, which plots the relationship between investor protection and income inequality.

4 Closed economy

In this section, I show briefly how the economy can be closed without affecting the main results discussed so far. Details of the analysis are provided in the appendix. Assume that capital and intermediate goods can no longer be imported or exported. Therefore, their prices will be pinned down by domestic demand and supply: $r_t = \alpha \frac{Y_t}{K_{Yt}}$, and $\chi_t = (1-\alpha)\frac{Y_t}{X_t}$. Further, capital will follow the law of motion:

$$\begin{aligned} K_{t+1} &= \frac{1}{1+\beta}\left\{G\left(\pi_t^*\right)B\chi_t + A\chi_t\int_{\pi_t^*}^{1}\left[\pi+(1-\pi)\varphi\right]g\left(\pi\right)d\pi \right. \\ &\quad \left. -\left[1-G\left(\pi_t^*\right)\right]r_t\right\}, \end{aligned} \tag{2.4}$$

where the RHS is aggregate savings. Aggregate capital is allocated between the final good sector and risky activities:

$$K_{t+1} \equiv K_{Yt+1} + 1 - G\left(\pi_{t+1}^*\right).$$

The aggregate supply of intermediate goods, X_t, equals total production of safe and risky projects:

$$X_t = G\left(\pi_t^*\right)B + A\int_{\pi_t^*}^{1}\left[\pi+(1-\pi)\varphi\right]g\left(\pi\right)d\pi.$$

Optimal technology adoption maintains the threshold property of Lemma 1, since agents take prices as given and the risky payoffs are still increasing in ability. In any period, the threshold ability π_t^* satisfies:

$$\pi_t^* v\left(w_t^h\left(\pi_t^*\right), r_{t+1}\right) + \left(1-\pi_t^*\right)v\left(w_t^l\left(\pi_t^*\right), r_{t+1}\right) = v\left(B\chi_t, r_{t+1}\right). \tag{2.5}$$

Differently from the small open economy, equilibrium payoffs $w_t\left(\pi_i\right)$ now depend also on the capital used in the final sector, K_{Yt}.

Equations (2.5) and (2.4) characterize the dynamic equilibrium. In the appendix, I report numerical solutions for the steady state and the transition dynamics. In particular, I show that Lemmas 2-3 and Propositions 1-2 continue to hold in the steady state. Moreover, along the transition between steady states with different investor protection, stock market size converges monotonically. Income inequality may instead converge along an oscillatory path, as a consequence of the dynamics of prices and capital.

5 Empirical evidence

The model developed through sections 2 and 3 generates three main results. (1) Stock markets are more developed, the better is investor protection. (2) Income inequality has an inverted-U shaped relationship with stock market size, both in (a) absolute terms and (b) relative to GDP. (3) Investor protection only affects income inequality through stock market size. Here, I empirically assess all the results by applying a series of cross-section and panel data methodologies. The section is structured as follows: I first present the cross-sectional and panel data techniques to be used, then the data, and finally report and comment on all the results.

5.1 Estimation strategies

5.1.1 Cross-section

To test the predictions of the model, I estimate the following static equation:

$$g_{i(t-k,t)} = \alpha + \beta \mathbf{x}_{i(t-k,t)} + \gamma_1 smdev_{i(t-k,t)} + \gamma_2 \left(smdev_{i(t-k,t)}\right)^2 + \epsilon_i, \tag{2.6}$$

where $g_{i(t-k,t)}$ is the Gini coefficient observed in country i over the period between $t-k$ and t, the terms in $\mathbf{x}_{i(t-k,t)}$ are additional explanatory variables, and $smdev_{i(t-k,t)}$ is the measure of stock market development. All variables are expressed in logaritm. To test both versions of result (2), I use two proxies for $smdev$: the ratios of stock market capitalization over GDP ($smcap$) and over credit to the private sector ($smpr$). The second variable measures the weight of equity finance over the total external finance (broadly, equity plus debt). It has also been used in the literature to proxy the overall degree of risk-sharing that can be achieved through the financial market. I select the regressors in $\mathbf{x}_{i(t-k,t)}$ so as to match the technology and skill parameters of the model with observable counterparts, and to control for factors commonly given attention in the empirical literature on inequality. $\mathbf{x}_{i(t-k,t)}$ includes time $t-k$ GDP and GDP squared to account for technology and the Kuznets hypothesis. I take two measures of the initial education attainment to proxy both the level and the dispersion of human capital. In particular, I use the share of the population aged above 25, with some secondary education (sec 25), and the Gini coefficient for the years of education in the population aged above 15 (gh_15). I control for government expenditure and trade openness to check the robustness of the results, and replace $g_{i(t-k,t)}$ with g_{it} for sensitivity analysis. Result (2) is confirmed by the data if $\hat{\gamma}_1 > 0$ and $\hat{\gamma}_2 < 0$. Notice, however, that g in the model may start to decline with *smdev* at high levels of stock market development that are rarely observed. As a consequence, the significance of $\hat{\gamma}_2$ might be weak in the data.

Equation (2.6) only captures the main result (2) of the paper (the inverted-U shaped relationship between stock market development and income inequality). To account for the intermediate link between investor protection and the size of the stock market (results (1) and (3)), I also estimate equation (2.6) by Two-Stages Least Squares, using a number of investor protection indicators as instruments for $smdev_{i(t-k,t)}$:

$$\begin{aligned} g_{i(t-k,t)} &= \alpha + \beta \mathbf{x}_{i(t-k,t)} + \gamma_1 smdev_{i(t-k,t)} + \gamma_2 \left(smdev_{i(t-k,t)}\right)^2 + e_i \\ smdev_{i(t-k,t)} &= \zeta + \xi \mathbf{ip}_{i(t-k,t)} + u_i. \end{aligned}$$

I adopt two alternative sets of instruments, $\mathbf{ip}_{i(t-k,t)}$, for stock market development following the analysis in La Porta et al. (LLS, 2003): (i) the indicators of investor protection and efficiency of the judiciary suggested by LLS as determinants of stock market development; (ii) the origin of the legal system which is, in turn, used by LLS to instrument investor protection. The advantages of the second set of instruments are that these are most certainly exogenous and available for a wider cross-section of countries. The IV estimation validates result (1), if $\hat{\xi} > 0$ and the F statistics of the excluded instruments from the first-stage regression is high. Result (3) is supported by the data, if the Sargan test of overidentifying restrictions has a high p-value, excluding correlation between investor protection and the residuals $e_{i.}$

5.1.2 Fixed and random effects

To test if the results of the paper hold both across countries and over time, I use the panel data methodology and estimate the following equation:

$$g_{it} = \alpha + \beta' \mathbf{x}_{it} + \gamma_1 smdev_{it} + \gamma_2 \left(smdev_{it}\right)^2 + \eta_i + \nu_t + \epsilon_{it}, \tag{2.7}$$

where g_{it} is the Gini coefficient observed in country i over a five-year period t, the terms in $\mathbf{x}_{it}$ and $smdev_{it}$ are the same as for equation (2.6), and η_t, ν_t and ϵ_{it} are unobservable country- and time-specific effects, and the error term, respectively. I estimate equation (2.7) under the alternative hypotheses of a random versus fixed idiosyncratic component η_i. Fixed-effects estimates capture the evolution of the relationship within each country over time. Random effects are more efficient, since they exploit all the information available across countries and over time. However, the latter may be inconsistent if country-specific effects are correlated with the residuals. Including time fixed effects in both regressions allows me to account for the presence of trends, such as skill-biased technical change, which drives inequality worldwide. I rely on the Hausman test for the choice between FE and RE, and an F test for the inclusion of time dummies.

5.1.3 Dynamic Panel Data

As a further evaluation of result (2) in a dynamic setting, I follow the approach of the latest studies on growth and inequality, and focus on the expression:

$$g_{it} = \lambda g_{it-1} + \tilde{\beta}' \mathbf{x}_{it} + \tilde{\gamma}_1 smdev_{it} + \tilde{\gamma}_2 \left(smdev_{it}\right)^2 + \eta_i + \nu_t + \epsilon_{it}. \tag{2.8}$$

Notice that the specification in equation (2.8) includes a lagged endogenous variable among the regressors. It immediately follows that, even if ϵ_{it} is not correlated with g_{it-1}, the estimates are not consistent with a finite time span. Moreover, consistency may be undermined by the endogeneity of other explanatory variables, such as income and stock market development. A number of contributions provide theoretical support (see, for instance, Banerjee and Duflo, 2003, Barro, 2000, Benabou, 1997, Forbes, 2003, and Lopez, 2003) and empirical treatments for the simultaneity between growth and inequality. Feedbacks with stock market size instead capture the reaction of capital supply to changes in the income distribution. To correct for the bias created by lagged endogenous variables, and the simultaneity of some regressors, I adopt the approach of Arellano and Bover (1995) and Blundell and Bond (1998).[18] I time-differentiate both sides of (2.8) to obtain

$$\Delta g_{it} = \lambda \Delta g_{it-1} + \tilde{\beta}' \Delta \mathbf{x}_{it} + \tilde{\gamma}_1 \Delta smdev_{it} + \tilde{\gamma}_2 \Delta \left(smdev_{it}\right)^2 + \Delta\nu_t + \Delta\epsilon_{it}, \tag{2.9}$$

and estimate the system of equations (2.8) and (2.9). The differences in the variables that are either endogenous or predetermined can be instrumented with their

[18]The system-DPD methodology dominates the difference-DPD proposed by Arellano and Bond (1991), because it amends problems of measurement error bias and weak instruments, arising from the persistence of the regressors (as pointed out by Bond et al., 2001).

own lagged values, while lagged differences are instruments for levels. For instance, I use g_{it-3} as an instrument for Δg_{it-1} and $smdev_{it-2}$ for $\Delta smdev_{it}$, as well as Δg_{it-2} and $\Delta smdev_{it-1}$ for g_{it-1} and $smdev_{it}$. The estimation is performed with a two-step System-GMM technique. The moment conditions for the equation in differences are $E[\Delta g_{it-s}\ (\epsilon_{it} - \epsilon_{it-1})] = 0$ for $s \geq 2$, and – if the explanatory variables y are predetermined – $E[\Delta y_{it-s}\ (\epsilon_{it} - \epsilon_{it-1})] = 0$ for $s \geq 2$. For equation (2.8), the additional conditions are $E[\Delta g_{i,t-s}\ (\eta_i + \varepsilon_{i,t})] = 0$ and $E[\Delta y_{i,t-s}\ (\eta_i + \varepsilon_{i,t})] = 0$ for $s = 1$. The validity of the instruments is guaranteed under the hypothesis that ϵ_{it} exhibit zero second-order serial correlation. Coefficient estimates are consistent and efficient, if both the moment conditions and the no-serial correlation are satisfied. I can validate the estimated model through a Sargan test of overidentifying restrictions, and a test of second-order serial correlation of the residuals, respectively. As pointed out by Arellano and Bond (1991), the estimates from the first step are more efficient, while the test statistics from the second step are more robust. Therefore, I will report coefficients and statistics from the first and second step, respectively.

5.2 Data

I use two cross-sections and two unbalanced panel datasets. The cross-section includes observations for 69 countries for the period 1980-2000. The sample shrinks to 43 observations when I account for investor protection and efficiency of the judiciary in the regressions, since these variables are only available for 49 countries, some of which do not intersect with the wider dataset. The main panel consists of 157 non-overlapping five-year observations, at least two for each of 52 countries, over the period 1976-2000. Since 16 countries have less than the three subsequent observations needed for the Arellano and Bover (1995) estimation, I use the full dataset only for the static panel regressions. I perform the dynamic panel GMM, as well as further static regressions, on a restricted sample of 125 observations for 36 countries over the same time span.

As a measure of income inequality, I take the Gini coefficients from Dollar and Kraay's (2002) database on inequality which relies on four sources: the UN-WIDER World Income Inequality Database, the "high quality" sample from Deininger and Squire (1996), Chen and Ravallion (2001), and Lundberg and Squire (2000).[19]

[19] The original sample consists of 953 observations, which reduce to 418 separated by at least five years, on 137 countries over the period 1950-1999. Countries differ with respect to the survey coverage (national vs subnational), the welfare measure (income vs expenditure), the measure of income (net vs gross) and the unit of observation (households vs individuals). Data from Deininger

Data on stock market capitalization (*smcap*) as a ratio of GDP and credit to the private sector (*privo*) on GDP come from the database of Beck et al. on Financial Development and Structure, which expands the data used in Beck et al. (1999). Their ratio is $smpr \equiv \frac{smcap}{privo}$.

The series for real per capita GDP, government expenditure and trade as a share of GDP are taken from Heston and Summers' version 6.1 of the Penn World Tables. Throughout the estimations, real per capita GDP is expressed as a ratio of the first observation for US GDP (1980 in the cross-section, 1976 in the panel).

I use two measures of human capital. The first is the percentage of people older than 25 years who have completed or are enrolled in secondary education ($sec25$). Data are taken from Barro and Lee's (2000) database. The second measure, better suited to capture the distribution of human capital, is the Gini coefficient of school years (gh_15) constructed by Castellò and Doménech (2002) on data from Barro and Lee (2000).

The indicators of investor protection and efficiency of the judiciary come from LLS(2003). Both $investor\_pr$ and eff_jud are indexes scaling from 0 to 10 in ascending order of protection and efficiency. See LLS (2003) for a detailed description.

The data on legal origins are taken from the World Development Indicators.

5.3 Results

5.3.1 Cross-sectional regressions

Table 2 reports the Ordinary Least Squares estimates for different versions of equation (2.6). Columns 1-10 suggest human capital and stock market development to be the major forces driving income inequality over the sample of 69 countries. As predicted by the model, $\hat{\gamma}_1$ is positive and significant for both stock market capitalization and its ratio to private credit, while $\hat{\gamma}_2$ is negative, though only significant for $smpr$. Notice that, according to these estimates, stock market development should start reducing inequality after reaching levels so high that five countries at most would be on the declining part of the $Gini\,(smcap)$ schedule, and nine in the case of $Gini\,(smpr)$. Thus, it seems that only very few countries have reached the point where the relationship between stock market size and inequality becomes negative. This may explain the low statistical significance for $\hat{\gamma}_2$. Moreover, the model predicts

and Squire are usually adjusted by adding 6.6 to the Gini coefficients based on expenditure. Here, the adjustment was made in a slightly more complicated way to account for the variety of sources; see Dollar and Kraay (2002) for details.

that inequality should never completely revert, even when the stock market achieves its maximum development; hence, it is reasonable to expect the linear term to be generally more relevant, as is the case in Table 2.

The significantly negative coefficients on sec 25 through columns 1-4 and 9-10, in line with most empirical evidence, mean that inequality tends to be lower, the larger is the share of the population with high education. The positive and significant estimates for gh_15 in columns 5-8 show that the dispersion of human capital boosts income inequality. However, the coefficients for sec 25 and gh_15 jointly estimated (Columns 9-10) suggest that the former is more effective at reducing inequality than the latter is at raising it. Given that sec 25 dominates gh_15, I will henceforth report the results obtained with sec 25 only. Finally, for the Kuznets hypothesis to hold, the estimated coefficients of GDP and $(GDP)^2$ should be positive and negative, respectively. The results in Table 2 do not allow me to validate this hypothesis, due to the lack of significance of both coefficients.

To get a quantitative flavor of the implications of columns 2 and 4, take pairs of countries with similar human capital (the other main determinant of inequality) but different stock market development, and compare the actual Gini differentials with their predicted values. Ecuador and Bolivia, for instance, had roughly the same school attainment (23.6 and 23 per cent of the population aged above 25 with secondary education), while stock market capitalization over GDP was 2.5 times larger in Ecuador. Column 2 would predict a lower Gini coefficient in Bolivia, with a three per cent difference: very close to the actual 3.1 per cent. Consider also Austria, which had the same level of secondary school attainment as Switzerland (65.1 vs 65.3), but a much less developed stock market (*smpr* was seven times smaller). Its predicted Gini (from the estimates in column 4) is lower than the Swiss by 19.1 vs the actual 19.7 per cent.

The results in Table 2 support the main prediction of the model on the relationship between stock market development and income inequality, but cannot provide evidence on the mechanism generating it, starting from investor protection. To ascertain that investor protection does not affect income inequality unless through stock market development, I first regress the Gini coefficient on the control variables in $\mathbf{x}$ and LLS's indicator of investor protection, and then add *smdev*. Table 3 shows that $investor\_pr$ indeed has a positive and significant effect on income inequality. However, the coefficients in columns 2 and 3 suggest that this effect is absorbed by stock market development, once controlled for. Moreover, columns 3 and 5 support the hypothesis that investor protection has no effect on inequality, unless paired by a thicker stock market. These results suggest that investor protection only affects income inequality through the development of equity markets.

The instrumental variables estimates reported in Tables 4 and 5 are meant to explicitly account for the intermediate step linking stock market development to the degree of investor protection. Estimating the first step of the IV regressions allows me to partially replicate the analysis in LLS (2003) to verify the predictive power of investor protection and efficiency of the judiciary on both indicators of stock market development. The coefficients in columns 1 and 4 of Table 5 confirm that better contractual protection boost stock market development, relative to the size of the economy and the overall financial depth. Since these variables can be suspected to be endogenous, I replace them with legal origins when estimating the first stage for *smcap* and *smpr*. Columns 1 and 3 confirm the results in LLS (2003) that the common law (UK) legal origin strongly promoted the development of stock markets. The results from only including the instrumented linear term of *smdev* in the regression for the Gini's (odd columns of Table 4) strongly support the prediction $\gamma_1 > 0$. P-values of the F and Sargan tests guarantee that both sets of instruments are valid. In other words, investor protection is a good predictor for *smdev* (result 1), but only affects inequality through stock market development (result 3). Estimating the equation with both linear and quadratic instrumented *smdev*, delivers a worse fit and insignificant coefficients for almost all covariates (also sec 25 loses significance in one case). However, the coefficient estimates from the first step suggest the existence of collinearity between the two sets of instruments, which undermines the validity of this specification.

So far, I have regressed average Gini coefficients on average stock market development. To verify if the results are sensitive to the timing of observations, the estimates of Tables 2 and 4 are replicated in two alternative ways. First, I replace the average Gini with its latest available observation and keep the regressors as in the previous estimates. The results are reported in Table 6. As a further check, I focus on the period 1985-2000 and regress the average Gini on the initial values of *smcap* and *smpr*. In this case, I do not need to perform instrumental variables estimations. As shown by Table 6b, one third of the observations gets lost. This can partly motivate the insignificance of γ_2, since a relevant part of the countries on the right-hand side of the hump is missing. Overall, this evidence favors the existence of a positive γ_1, and a weaker negative γ_2.

Finally, the robustness of the results is tested in Table 7, which reports the estimates of equation (2.7) where government expenditure and trade (as a ratio of

GDP) are added as additional covariates. There are no major changes from Tables 2 and 4, and the additional coefficients are not significantly different from zero.

5.3.2 Panel regressions

Columns 1 and 2 in Table 8 report the coefficients of equation (2.7) estimated with fixed and random effects, respectively. Stock market development significantly affects income inequality, following a humped-shaped relation as predicted by the model. Four observations lay on the downward sloping part of the hump: Hong Kong and Malaysia in the period 1991-2000. When I control for time fixed effects, the significance of the quadratic term in stock market capitalization is weakened, while the positive linear relationship remains strong, as shown in columns 3 and 4. Education turns out to be negatively related to inequality throughout all estimations, consistently with most of the empirical literature. The Kuznets hypothesis is not validated by the results in Table 8. The results for the stock market as a ratio of private credit in the last two columns of Table 8 confirm the existence of a positive γ_1, but do not provide strong support for $\gamma_2 < 0$. In conclusion, the static panel analysis suggests that stock market development plays as important a role as education in shaping income distribution.

The regression in column 3 of Table 8 to some extent controls for the time variation in the relationship between changes in stock market development and income inequality within countries and across time. However, it does not account for the existence of dynamic feedbacks between inequality and stock market development. To overcome these methodological limitations, I adopt the approach of Arellano and Bover (1995) and Blundell and Bond (1998), and estimate various versions of system (2.8)-(2.9).

The results in Table 9 confirm the existence of a significant positive linear relationship between the Ginis and stock market development. The quadratic term is also significant and exhibits the expected negative sign, in the estimates for *smpr*. The positive γ_1 survives the inclusion of time, as well as time-continent effects.[20] All estimated coefficients for $d\log\left(Gini_{t-1}\right)$ support the convergence hypothesis for income inequality, as in previous empirical work by Benabou (1996), Lopez (2003) and Ravallion (2002). As in the previous evidence, the Kuznets' hypothesis finds no support and the effectiveness of human capital becomes weaker.

To make the results from dynamic and static panel regressions comparable, I

[20]Results with time-continent effects are available upon request.

replicated the Fixed and Random Effects estimates on the restricted sample and reported the coefficients in Table 10. The linear term for stock market development still has a positive and significant effect throughout all estimates, while the $\hat{\gamma}_2$ are non-significantly different from zero in all specifications.

As a robustness check, I re-estimate the equations in Tables 9-10 with government expenditure and trade over GDP as additional regressors. Table 11 reports the estimated coefficients for stock market capitalization, both in linear and non-linear terms, and for the new control variables. Both static and dynamic regressions support the prediction of a positive γ_1, while the negative γ_2 is only significant in the system-GMM for *smpr*. The estimates for government expenditure, which are non-significantly different from zero, reflect the ambiguity of theoretical predictions and previous empirical evidence. The coefficients for trade openness from Fixed Effects regressions point towards a positive effect on inequality, consistently with previous theoretical predictions and empirical evidence (see, for example, Epifani and Gancia, 2002, Feenstra and Gordon, 2001, Robbins, 1996, Spilimbergo et al., 1999). The dynamic panel data estimates support the opposite view, even though to a lesser extent, since the negative coefficients are significant at 8 per cent, at most.

5.3.3 Summary

The estimates reported in this section suggest that stock market development tends to raise income inequality. The declining part of the hump predicted by the model is supported in a less robust way by the data. This evidence can be reconciled with the model, since the peak of the Gini coefficient may only occur at such high levels of stock market development that are not observed in the sample. Dynamic Panel Data estimates suggest the relationship between stock market development and income inequality to hold in the long run, as predicted by the general equilibrium version of the model. Results from the cross-sectional regressions confirm the prediction that investor protection only affects income inequality through the development of the equity market.

6 Conclusions

This paper provides theoretical predictions and empirical support for a systematic relationship between investor protection, financial development and income inequality. I develop an overlapping generation model with risk-averse agents, heterogeneous in their ability, where production can take place with a safe or a risky technology. In the presence of financial frictions, arising from the non-observability of realizations and imperfect investor protection, I study the occupational and financial choices for different ability groups. Better investor protection promotes financial development and affects income inequality in a number of ways. First, it improves risk sharing, thereby reducing income volatility for a given size of the risky sector. Second, it raises the share of population exposed to earning risk. Finally, since ability affects risky payoffs, it increases the overall reward to ability. The first effect tends to reduce inequality, while the other two boost it. The main result of the paper is that income dispersion increases at first with financial development, and then declines. In the empirical section, I provide evidence consistent with the predictions of the model.

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7 Appendix

7.1 Proofs

Lemma 1

The assumptions that $A > B + r$ and $\varphi A < B + r$ together with continuity of V_i in π_i imply the existence of a unique point $\pi^* \in (0,1)$ where $V^* = B$. From this, it follows that for $\pi_i = 1$, $\left(1-\theta_i^h\right) A = (A-r) > Bx$, hence $V_i = v\left[\left(1-\theta_i^h\right) A, r\right] > v(B,r)$, and for $\pi_i = 0$, $\left(1-\theta_i^l\right) \varphi A = \varphi A - r < B$, thus $V_i = v\left[\left(1-\theta_i^l\right) \varphi A, r\right] < v(B,r)$. To prove that π^* is a threshold, I just need to show that V_i is increasing in π_i. The derivative of V_i w. r. t. π_i under the optimal equity contract is

$$\frac{dV_i}{d\pi_i} = v\left[\left(1-\theta_i^h\right) A, r\right] - v\left[\left(1-\theta_i^l\right) \varphi A, r\right] + \left[\pi_i v_h' + (1-\pi_i) v_l'\right] pA > 0.$$

Therefore, $\forall \pi_i \geq \pi^*$, $\pi_i v\left[\left(1-\theta_i^h\right) A, r\right] + (1-\pi_i) v\left[\left(1-\theta_i^l\right) \varphi A, r\right] \geq v(B,r)$.

Lemma 2

To prove that the threshold ability is decreasing in investor protection, I obtain the derivative of π^* with respect to p,

$$\frac{d\pi^*}{dp} = -\frac{dV}{dp}\left(\frac{dV}{d\pi}\right)^{-1},$$

and show that it is negative. I have derived $\frac{dV}{d\pi^*} > 0$ in the proof of Lemma 1. I just need to derive

$$\frac{dV}{dp} = \pi_i (1-\pi_i)(1-\varphi) A \left(v_l' - v_h'\right).$$

Notice that $\frac{dV}{dp} > 0$ for any π, since utility is concave. It follows that $\frac{d\pi^*}{dp} < 0$.

To prove that the threshold is convex in investor protection, I need to prove that $\frac{d^2\pi^*}{(dp)^2} > 0$.

$$\begin{aligned}
\frac{d^2\pi^*}{(dp)^2} &= \frac{\frac{d^2V}{d\pi dp}\frac{dV}{dp} - \frac{d^2V}{(dp)^2}\frac{dV}{d\pi}}{\left(\frac{dV}{d\pi}\right)^2} \\
&= -\left(\frac{dV}{d\pi}\right)^{-1}\left\{\pi^*(1-\pi^*) A \left(v_l' - v_h'\right) + A\left[\pi^* v_h' + (1-\pi^*) v_l'\right]\right. \\
&\quad \left. - pA^2\pi^*(1-\pi^*)(1-\varphi)\left(v_l'' - v_h''\right)\right\}\frac{d\pi^*}{dp} \\
&\quad -\left(\frac{dV}{d\pi}\right)^{-1}\left\{A^2(1-\varphi)^2\pi^*(1-\pi^*)\left[\pi^* v_h'' + (1-\pi^*) v_l''\right]\right\}.
\end{aligned}$$

All terms divided by $\frac{dV}{d\pi}$ are positive, since the CRRA specification of the utility function implies that $v_l' > v_h'$ and $v_l'' < v_h''$, and $\frac{d\pi^*}{dp} \leq 0$. Therefore, $\frac{d^2\pi^*}{(dp)^2} = -(>0)^{-1}\{(\geq 0) + (>0) - (\leq 0)\}(\leq 0) - (>0)^{-1}\{<0\} > 0$.

Proposition 1

To prove the increasing monotonicity of stock market size, and its concavity at high levels of investor protection, I derive

$$\begin{aligned}\frac{dsm}{dp} &= -g(\pi^*)\frac{d\pi^*}{dp}\\ \frac{d^2sm}{(dp)^2} &= -g'(\pi^*)\left(\frac{d\pi^*}{dp}\right)^2 - g(\pi^*)\frac{d^2\pi^*}{(dp)^2}.\end{aligned}$$

From Lemma 1, $\frac{d\pi^*}{dp} \leq 0$, that implies $\frac{dsm}{dp} \geq 0$; hence, the stock market size is increasing in investor protection. From Lemma 2, $\frac{d^2\pi^*}{(dp)^2} > 0$. Moreover, $\lim_{p\to 1}\frac{d\pi^*}{dp} = \lim_{p\to 1}\left(\frac{dV}{dp}\Big/\frac{dV}{d\pi}\right) = \lim_{p\to 1}\frac{\pi(1-\pi)(1-\varphi)\left[v'\left(w^l,r\right)-v'\left(w^h,r\right)\right]}{v\left(w^h,r\right)-v\left(w^l,r\right)+\left[\pi v'\left(w^h,r\right)+(1-\pi)v'\left(w^l,r\right)\right]pA} = 0$. It follows that sm is concave in p in a neighborhood of $p = 1$, since $\lim_{p\to 1}\frac{d^2sm}{(dp)^2} < 0$.

Corollary 1

Re-write $Y = C = \frac{1+r+\beta}{1+\beta}\left\{G(\pi^*)B + \int_{\pi^*}^{1}\{[\pi + (1-\pi)\varphi]A - r\}g(\pi)\,d\pi\right\}$. The first derivative of $\frac{sm}{Y}$ w.r.t. p is

$$\begin{aligned}\frac{d\frac{sm}{Y}}{dp} = &-\frac{d\pi^*}{dp}\frac{g(\pi^*)}{Y^2}\frac{1+r+\beta}{1+\beta}\left\{A\int_{\pi^*}^{1}[\pi + (1-\pi)\varphi]\,d\pi\right.\\ &\left.+B - A[1-G(\pi^*)][\pi^* + (1-\pi^*)\varphi]\right\}.\end{aligned}$$

Stock market as a ratio of GDP is increasing in investor protection, $\frac{d\frac{sm}{Y}}{dp} \geq 0$ for any $p \in [0,1]$, since $\frac{d\pi^*}{dp} \leq 0$ and the term in brackets is always positive. To prove concavity of $\frac{sm}{Y}$ in a neighborhood of $p = 1$, I derive

$$\begin{aligned}\frac{d^2\frac{sm}{Y}}{(dp)^2} &= -\frac{d^2\pi^*}{(dp)^2}\frac{g(\pi^*)}{Y^2}\Psi\frac{1+r+\beta}{1+\beta} - \left(\frac{d\pi^*}{dp}\right)^2\frac{1+r+\beta}{1+\beta}\left\{\frac{g'(\pi^*)}{Y^2}\Psi\right.\\ &\quad +2\frac{g(\pi^*)}{Y^3}\frac{dY}{d\pi^*}\Psi + \frac{g(\pi^*)}{Y^2}A[1-g(\pi^*)][\pi^* + (1-\pi^*)\varphi]\\ &\quad \left.+(1-\varphi)[1-G(\pi^*)]\right\},\\ \Psi &\equiv A\int_{\pi^*}^{1}[\pi + (1-\pi)\varphi]\,d\pi + B - A[1-G(\pi^*)][\pi^* + (1-\pi^*)\varphi].\end{aligned}$$

As $\lim_{p\to 1}\frac{d\pi^*}{dp} = 0$, while $\frac{d^2\pi^*}{(dp)^2} > 0$ at any p, $\lim_{p\to 1}\frac{d^2\frac{sm}{Y}}{(dp)^2} < 0$.

Lemma 3

To prove non monotonicity, I differentiate $Var(w)$ with respect to p:

$$\begin{aligned}
\frac{dVar(w)}{dp} &= \frac{d\pi^*}{dp}\left\{g(\pi^*)[B-E(w)]^2 - 2G(\pi^*)[B-E(w)]\frac{dE(w)}{d\pi^*}\right\} \\
&-\frac{d\pi^*}{dp}g(\pi^*)\left\{\pi^*\left[w^h(\pi^*)-E(w)\right]^2+(1-\pi^*)\left[w^l(\pi^*)-E(w)\right]^2\right\} \\
&+\frac{d\pi^*}{dp}\frac{dE(w)}{d\pi^*}2\int_{\pi^*}^1\left\{\pi\left[w^h-E(w)\right]+(1-\pi)\left[w^l-E(w)\right]\right\}g(\pi)\,d\pi \\
&+2\int_{\pi^*}^1\left\{\pi\frac{dw^h}{dp}\left[w^h-E(w)\right]+(1-\pi)\frac{dw^l}{dp}\left[w^l-E(w)\right]\right\}g(\pi)\,d\pi
\end{aligned}$$

$$\begin{aligned}
&= \frac{d\pi^*}{dp}g(\pi^*)\left\{[B-E(w)]^2-\pi^*\left[w^h(\pi^*)-E(w)\right]^2\right. \\
&\quad\left.-(1-\pi^*)\left[w^l(\pi^*)-E(w)\right]^2\right\} \\
&\quad-2(1-\varphi)A\int_{\pi^*}^1\pi(1-\pi)\left(w^h-w^l\right)g(\pi)\,d\pi.
\end{aligned}$$

Notice that the term in the first two lines represents the market size effect and is positive for all p, while the last line accounts for the risk sharing effect and is negative for all p.

For $p\to 0$, $\pi^*\to 1$, $E(w)\to B$, $w^h\to A-r$, $w^l\to\varphi A-r$. Therefore,

$$\lim_{p\to 0}\frac{dVar(w)}{dp} = -\frac{d\pi^*}{dp}g(1)(A-B-r)^2>0.$$

For $p\to 1$, $\pi^*\to\pi^*_{p=1}=\frac{(B-\varphi A)+r}{(1-\varphi)A}$, $w^h(\pi^*)-w^l(\pi^*)\to 0$, $w^h\left(\pi^*_{p=1}\right)\to w^l\left(\pi^*_{p=1}\right)=\left[\pi^*_{p=1}+(1-\pi^*_P)\varphi\right]A-r=B$, $\frac{d\pi^*}{dp}\to 0$. I study how $\frac{dVar(w)}{dp}$ approaches zero in a left neighborhood of $p=1$ by means of Taylor's first-order approximation. Notice that

$$\begin{aligned}
\frac{d^2 Var(w)}{(dp)^2} &= \left[\frac{d^2\pi^*}{(dp)^2} g(\pi^*) + \left(\frac{d\pi^*}{dp}\right)^2 g'(\pi^*)\right] \left\{[B - E(w)]^2 \right.\\
&\quad \left. -\pi^* \left[w^h(\pi^*) - E(w)\right]^2 - (1-\pi^*) \left[w^l(\pi^*) - E(w)\right]^2 \right\}\\
&\quad + \frac{d\pi^*}{dp} g(\pi^*) \left\{ 2 \frac{d\pi^*}{dp} \frac{dE(w)}{d\pi^*} \{[\pi^* + (1-\pi^*)\varphi] A - r - B\} \right.\\
&\quad + 2\pi^*(1-\pi^*)(1-\varphi)^2 A^2 - \frac{d\pi^*}{dp} \left\{ \left[w^h(\pi^*) - E(w)\right]^2 \right.\\
&\quad \left. \left. - \left[w^l(\pi^*) - E(w)\right]^2 \right\} + 2(1-\varphi)^2 A^2 \int_{\pi^*_{p=1}}^{1} \pi(1-\pi) g(\pi) d\pi.
\end{aligned}$$

It follows that, in a neighborhood to the left of $p = 1$,

$$\frac{dVar(w)}{dp} = 2(p-1)(1-\varphi)^2 A^2 \int_{\pi^*_{p=1}}^{1} \pi(1-\pi) g(\pi) d\pi < 0.$$

Proposition 2

Recall from Proposition 1 that sm is increasing in p. I characterize the relationship between stock market size and the variance of earnings by studying

$$\begin{aligned}
\frac{dVar(w)}{dsm} &= \frac{dVar(w)}{dp} \left(\frac{dsm}{dp}\right)^{-1}\\
&= -[B - E(w)]^2 + (1-\pi^*) \left[w^l(\pi^*) - E(w)\right]^2\\
&\quad + \pi^* \left[w^h(\pi^*) - E(w)\right]^2 + \left[\frac{d\pi^*}{dp} g(\pi^*)\right]^{-1} \times\\
&\quad 2(1-\varphi)^2 A^2 (1-p) \int_{\pi^*}^{1} \pi(1-\pi) g(\pi) d\pi
\end{aligned}$$

For $p \to 0$, $\pi^* \to 1$, $E(w) \to B$, $w^h \to A - r$, $w^l \to \varphi A - r$, hence

$$\lim_{p \to 0} \frac{dVar(w)}{dsm} = (A - B - r)^2 > 0.$$

For $p \to 1$, $\pi^* \to \pi^*_{p=1} = \frac{(B-\varphi A)+r}{(1-\varphi)A}$, $w^h(\pi^*) - w^l(\pi^*) \to 0$, $w^h\left(\pi^*_{p=1}\right) \to w^l\left(\pi^*_{p=1}\right) =$

$\left[\pi^*_{p=1} + \left(1-\pi^*_{p=1}\right)\right] A- r = B$, and $\frac{d\pi^*}{dp} \to 0$. It thus follows that

$$\begin{aligned}\lim_{p\to 1}\frac{dVar\left(w\right)}{dsm\left(p\right)} &= \lim_{p\to 1}\frac{\frac{d}{dp}\left[2\left(1-\varphi\right)^2 A^2\left(1-p\right)\int_{\pi^*}^1 \pi\left(1-\pi\right)g\left(\pi\right)d\pi\right]}{\frac{d}{dp}\left[\frac{d\pi^*}{dp}g\left(\pi^*\right)\right]}\\ &= 2\int_{\pi^*_{p=1}}^1 \pi\left(1-\pi\right)g\left(\pi\right)d\pi\frac{v\left(B\right)+Av'\left(B\right)}{\pi^*_{p=1}\left(1-\pi^*_{p=1}\right)g\left(\pi^*_{p=1}\right)v''\left(B\right)} < 0,\end{aligned}$$

since $v'' < 0$ for any CRRA utility function.

7.2 Closed economy

7.2.1 The dynamics

The dynamics of the closed economy satisfies equations (2.4) and (2.5):

$$\pi^*_t v\left(w^h_t\left(\pi^*_t\right), r_{t+1}\right) + \left(1-\pi^*_t\right)v\left(w^l_t\left(\pi^*_t\right), r_{t+1}\right) = v\left(B\chi_t, r_{t+1}\right)$$

$$\begin{aligned}K_{t+1} &= \frac{1}{1+\beta}\left\{G\left(\pi^*_t\right)B\chi_t + A\chi_t\int_{\pi^*_t}^1\left[\pi+\left(1-\pi\right)\varphi\right]g\left(\pi\right)d\pi\right.\\ &\quad \left.-\left[1-G\left(\pi^*_t\right)\right]r_t\right\}.\end{aligned}$$

As noticed in section 4, earnings depend on factor prices, which are functions of π^*_t and capital employed in the final sector, $K_{Yt} = K_t - 1 + G\left(\pi^*_t\right)$. This implies that the threshold ability π^*_t becomes an implicit function of K_t and the analytical characterization of the dynamic equilibrium becomes awkward. Therefore, I proceed by means of numerical solutions. The main results are displayed in Figures 2.4-2.5. In all simulations, I adopt the following parametrization: $A = 120$, $B = 100$, $\alpha = 0.33$, $\beta = 0.17$ (equivalent to a six per cent annual discount for thirty years, i.e. a generation), and G uniform in [0, 1].

Figure 2.4 describes the dynamics of an economy that starts with a very low capital endowment, K_0, and an intermediate degree of investor protection, $p = 0.5$. When K_0 is too low ($K_0 < \alpha B/(1-\alpha)(A-B)$), the interest rate is so high relative to the price of the intermediate good that no young agent chooses the risky technology. Hence, there is no stock market and inequality is zero. As capital is accumulated, the interest rate falls and the price of intermediates rises. When the ratio r/χ becomes low enough, some young agents prefer the risky project and raise capital through equities. This requires a shift of capital out of the final good sector, which in turn tends to raise r and lower χ. As a result, with capital accumulation and an expanding stock market, r/χ falls by less than it would in the absence of the risky technology. Also, a positive stock market size implies that some income inequality arises due to the "market size" effect, as in the model of sections 2-3. Moreover, the ratio between factor prices, r/χ, also affects inequality by changing the earnings differentials between safe and risky entrepreneurs. The lower the ratio, the wider the earnings differentials, the higher inequality ("relative factor prices" effect). This implies that, with endogenous prices, inequality may vary even if stock market size does not. The adjustment of capital and prices continues until the steady state is reached. Decreasing marginal productivity of capital guarantees the existence of the steady state.

Figure 2.5 shows the adjustment after a policy change that increases investor protection from $p = 0$ to $p = 0.05$, starting from the steady state. Due to the convexity of π_t^* in p, the risky intermediate sector expands remarkably in response to the policy change. The marginal productivity of capital rises sharply both because some capital is shifted to the risky sector and because the production of intermediates increases. This causes an overshooting of the interest rate, that gradually declines with capital accumulation to its new (higher) steady state level. Inequality immediately jumps up and oscillates around its new (higher) steady state level until capital and prices are stable.

If the policy change occurs at high levels of investor protection, the effect on productivity of factors (hence prices) is weaker. An increase in p induces a small shift of capital from the final to the risky intermediate sector, and has almost no effect on the interest rate. Inequality falls, since the "risk sharing" effect outweighs the "market size" effect at high levels of investor protection.

7.2.2 The steady state

In the steady state, $K_{t+1} = K_t = K$ and $\pi_{t+1}^* = \pi_t^* = \pi^*$. The equilibrium is the solution to the system:

$$\begin{aligned} VV &\equiv \pi^* v\left(w^h(\pi^*), r\right) + (1-\pi^*)\, v\left(w^l(\pi^*), r\right) - v(B\chi, r) = 0 \\ KK &\equiv (1+\beta) K - G(\pi^*) B\chi - \int_{\pi^*}^{1} \left[\pi w^h(\pi) + (1-\pi)\, w^l(\pi)\right] g(\pi)\, d\pi = 0. \end{aligned}$$

The risky intermediate sector is active, at least in the presence of perfect investor protection, provided that $A - B > \left(\frac{1+\beta}{1-\alpha}\right)^{\frac{1}{1-\alpha}} \frac{\alpha}{1-\alpha}$. Comparative statics for p in the

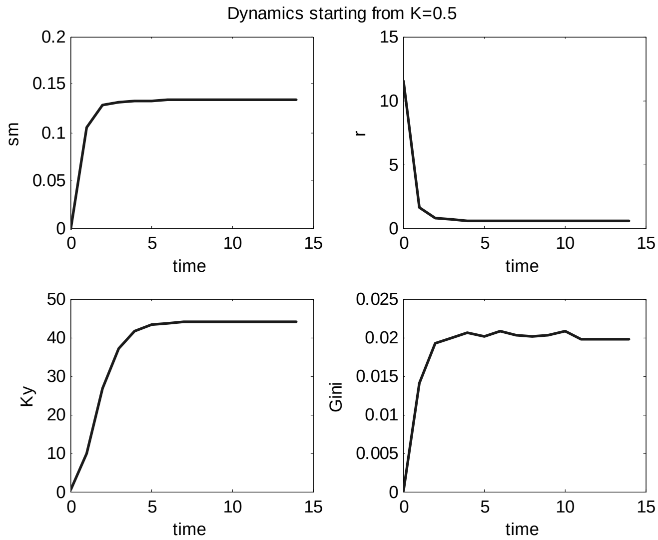

Figure 2.4: Dynamics from a low initial capital endowment (K=*0.5*) to the steady state, given p=*0.5*.

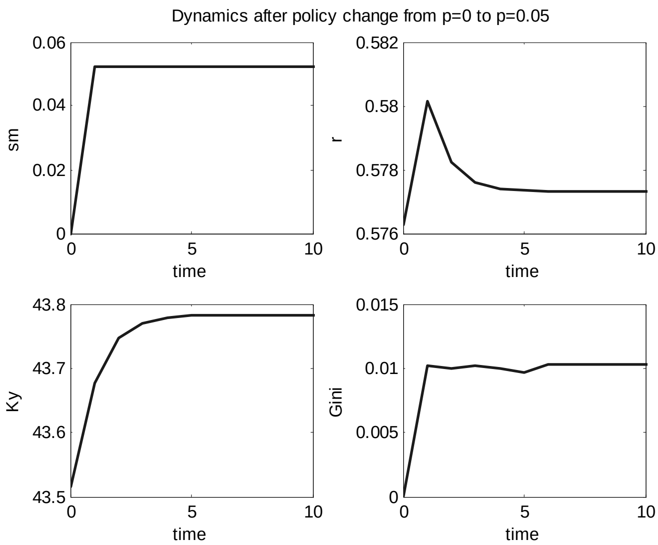

Figure 2.5: Dynamic adjustment after a policy change from p=*0* to p=*0.05*.

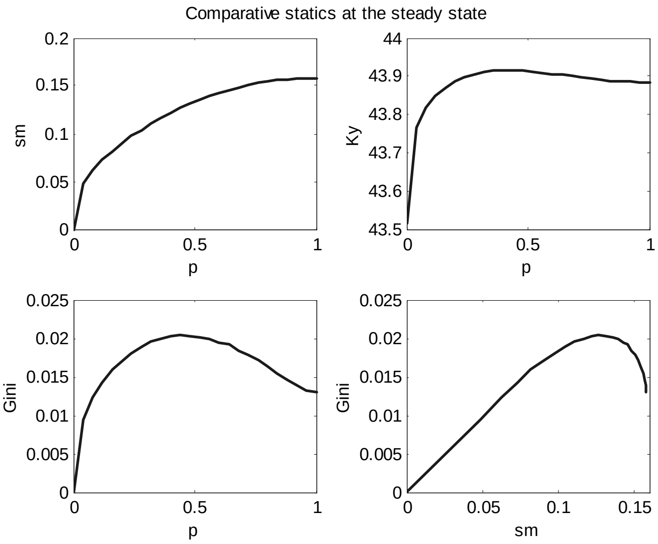

Figure 2.6: Comparative statics for investor protection in the steady state.

steady state are depicted in Figure 2.6 showing that Lemmas 1-3 and Propositions 1-2 continue to hold in the closed economy. In fact, the "relative factor prices" effect, that affects inequality along the dynamics, is irrelevant in the steady state. Therefore, the comparative statics on investor protection is driven by the "market size" and "risk sharing" effects only, as in the small open economy.

7.3 Simulation details

This section describes the procedure I followed for simulating the small open economy of sections 2-3 step by step.

1. Give values for the main parameters (A, B, φ, β, α) and the interest rate, and compute the threshold ability with perfect investor protection $\left(\pi^*_{p=1}\right)$.

2. Compute values for the parameters of the Lognormal distribution of abilities, (μ, σ), from Barro and Lee's (2000) data. The database provides observations for the percentages of the population aged 15 and above with no, primary, secondary and tertiary education (lu, lp, ls, lh), along with the average year of each education level (pyr, syr, hyr). I compute the average years of schooling for people with primary, secondary and tertiary education (q_1, q_2, q_3, respec-

tively):

$$q_1 = \frac{pyr}{lp + ls + lh}; q_2 = q_1 + \frac{syr}{ls + lh}; q_3 = q_1 + q_2 + \frac{hyr}{lh}.$$

The average years of schooling and their variance are then

$$\begin{aligned} E(Q) &= \sum_{i=1}^{3} l_i q_i \\ V(Q) &= \sum_{i=0}^{3} l_i \left(q_i - E(Q)\right)^2, \end{aligned}$$

with $l_0 = lu$, $l_1 = lp$, $l_2 = ls$ and $l_3 = lh$. Group the countries in low-income, middle-income and high-income following the WDI criterion and take the average values of $E(Q)$ and $V(Q)$. Finally, μ and σ can be derived from the expressions for mean and variance of the Lognormal distribution:

$$\begin{aligned} E(Q) &= e^{\mu + \frac{\sigma^2}{2}} \\ V(Q) &= e^{2\mu + 2\sigma^2} - e^{\mu + \sigma^2}. \end{aligned}$$

3. Define a grid of 101 degrees of investor protection $p \in [0, 1]$, and a grid of initial guesses for the threshold ability $\pi^* \in \left[\pi^*_{p=1}, 1\right]$, equally spaced by 0.0001 (the finer the grid, the better the approximation).

4. Draw Π =10001 ability levels from a Lognormal (μ, σ) and sort them in ascending order. Identify the ability level $\pi_{.9995}$: $G(\pi_{.9995}) = 0.9995$ and divide every $\pi \leq \pi_{.9995}$ by this figure. Replace all $\pi > \pi_{.9995}$ by 1, so that the distribution is normalized to values included in $[0, 1]$, and truncated in a way that makes the top 0.05 per cent of the population successful with certainty. Compute the Cdf of ability,

$$G(\pi_i) = \frac{\#\text{ of realizations } \pi \leq \pi_i}{\Pi}.$$

5. For every degree of investor protection p

   (a) compute $\pi^*(p)$ as the solution to the technology choice problem. In particular, recursively find the point in the grid of π^* satisfying:

$$\begin{aligned} \log(B) &= \pi^* \log\left(w^h\right) + (1-\pi^*)\log\left(w^l\right) \\ w^h &= A\left[\pi^* p(1-\varphi) + \varphi + (1-p)(1-\varphi)\right] - r \\ w^l &= A\left[\pi^* p(1-\varphi) + \varphi\right] - r > 0. \end{aligned} \tag{2.10}$$

(b) For every ability π

i. draw the earning realization:

$$\begin{aligned} w &= \begin{cases} B & \pi < \pi^* \\ A\left[\pi^* p(1-\varphi) + \varphi + (1-p)(1-\varphi)\epsilon\right] - r & \pi \geq \pi^* \end{cases} \\ \epsilon &\sim Bi(N,\pi), \text{ with } N = \#\text{ of } \pi \geq \pi^*. \end{aligned}$$

ii. sort w and derive its cumulative density function as $F(w_i) = \frac{\#\text{ of realizations } w \leq w_i}{\Pi}$

iii. compute the Lorenz Curve as $L(w_m) = \frac{\textit{mean of } w \leq w_m}{\textit{mean of } w}\frac{m}{\Pi}$ for $m = 1, 2, ...\Pi$

iv. compute the Gini coefficient as $Gini = 1 - 2\sum_{m=1}^{\Pi} \frac{L(w_m)}{\Pi}$

(c) save the threshold and the Gini in $(1 \times p)$ vectors, $\boldsymbol{\pi}^*(p)$ and $Gini(p)$, the earnings realizations, their distribution and the Lorenz curve in $(p \times \Pi)$ matrices, $\mathbf{w}(p,\pi)$, $\mathbf{F}(p, w(p,\pi))$ and $\mathbf{L}(p, w(p,\pi))$

When simulating the closed economy, step 1 does not specify r.

Step 5.(a) finds the threshold ability $\pi_t^*(p)$ which solves (2.10) for a given initial capital K_t, taking into account that $\chi_t = (1-\alpha)\left[K_t - 1 + G(\pi_t^*)\right]^{\alpha}\left\{A\sum_{i=\pi_t^*}^{1}\left[\pi_i + (1-\pi_i)\varphi\right] g(\pi_i) + G(\pi^*)B\right\}^{-\alpha}$ and $r_t = \alpha\left[K_t - 1 + G(\pi_t^*)\right]^{\alpha-1}\left\{A\sum_{i=\pi_t^*}^{1}\left[\pi_i + (1-\pi_i)\varphi\right] g(\pi_i) + G(\pi^*)B\right\}^{1-\alpha}$.

After step 5.(c), capital in the next period is computed as $K_{t+1} = \sum_{i=0}^{1} w_i - \left[1 - G(\pi_t^*)\right] r$ and plugged into step 5.a. as new initial capital K_t. This recursion goes on until the steady state is reached and $K_t = K_{t+1}$.

Table A
Countries and Samples

Country	CL	CS	PL	PS	Country	CL	CS	PL	PS
Australia	y	y	y	y	Kenya	y	y		
Austria	y				Korea	y	y	y	y
Bangladesh	y		y	y	Malaysia	y	y	y	y
Barbados	y				Mauritius	y		y	
Belgium	y	y	y		Mexico	y	y	y	y
Bolivia	y				Nepal	y			
Botswana	y				Netherlands	y	y	y	y
Bulgaria			y		New Zealand	y	y	y	y
Brazil	y	y	y	y	Norway	y	y	y	y
Canada	y	y	y	y	Pakistan	y	y	y	y
Chile	y	y	y		Panama	y			
China	y		y		Paraguay	y			
Colombia	y	y	y		Peru	y	y	y	y
Costa Rica	y		y	y	Philippines	y	y	y	
Denmark	y	y	y	y	Poland	y		y	y
Ecuador	y	y	y		Portugal	y	y	y	
Egypt	y	y	y		Romania	y			
El Salvador	y				Russia			y	y
Finland	y	y	y	y	Singapore	y	y	y	y
France	y	y	y	y	Slovak Republic			y	
Germany	y	y	y	y	South Africa	y	y		
Ghana	y		y	y	Spain	y	y	y	y
Greece	y		y		Sri Lanka	y	y	y	y
Guatemala	y				Sweden	y	y	y	y
Honduras	y				Switzerland	y			
Hong Kong	y	y	y	y	Taiwan	y	y	y	y
Hungary	y		y		Thailand	y	y	y	y
India	y	y	y	y	Trinidad and Tobago	y		y	y
Indonesia	y	y	y	y	Tunisia	y		y	
Iran	y				Turkey	y	y	y	
Ireland	y				United Kingdom	y	y	y	y
Israel	y	y			United States	y	y	y	y
Italy	y	y	y	y	Uruguay	y	y		
Jamaica	y		y		Venezuela	y	y	y	y
Japan	y	y	y	y	Zambia	y			
Jordan	y	y	y	y	Zimbabwe	y	y		

Note: C and P stand for cross-sectional and panel datasets, respectively.
L and S for large and small samples.

Table 2. Stock market development and income inequality

cross-section - 1980-2000

	OLS	OLS	OLS	OLS	OLS	OLS	OLS	OLS	OLS	OLS
smcap	**.110** (.030)	**.192** (.085)			**.098** (.029)	.125 (.09)			**.104** (.033)	
$smcap^2$		−.093 (.097)				−.031 (.095)				
smpr			**.073** (.033)	**.179** (.055)			**.067** (.034)	**.186** (.059)		**.068** (.034)
$smpr^2$				*−.083* (.048)				*−.091* (.049)		
sec 25	**−.214** (.057)	**−.233** (.059)	**−.191** (.053)	**−.199** (.052)					*−.156* (.091)	−.128 (.092)
gh_15					**.158** (.057)	**.164** (.058)	**.151** (.06)	**.179** (.064)	.052 (.089)	.063 (.095)
GDP	−.159 (.128)	−.169 (.128)	−.053 (.141)	−.085 (.139)	−.075 (.134)	−.073 (.135)	.054 (.145)	.013 (.143)	−.137 (.136)	−.025 (.150)
GDP^2	.175 (.183)	.188 (.187)	.084 (.204)	.117 (.201)	−.022 (.18)	−.025 (.181)	−.126 (.198)	−.079 (.192)	.124 (.198)	.026 (.216)
R^2	.550	554	.546	575	.519	.520	.521	.554	.543	.538
Obs	69	69	68	68	68	68	67	67	68	67

The dependent variable is the average Gini coefficient between 1980 and 2000. GDP and education are the initial values, stock market development is the average. All regressions include a dummy for Latin America. Coefficients are estimated with Ordinary Least Squares. Robust standard errors within parenthesis, 5% and 10% significant coefficients in bold and italics, respectively.

Table 3. Stock market size, investor protection and income inequality

cross-section - 1980-2000

	OLS	OLS	OLS	OLS	OLS
investor_pr	**.006** (.003)	.003 (.003)	−.001 (.004)	−.0001 (.004)	*−.007* (.004)
smcap		*.070* (.042)			
smpr				**.121** (.041)	
smcap ∗ investor_pr			*.012* (.064)		
smpr ∗ investor_pr					**.018** (.005)
sec 25	**−.174** (.065)	**−.156** (.061)	**−.141** (.059)	**−.174** (.063)	**−.145** (.057)
GDP	−.086 (.258)	.035 (.128)	.031 (.121)	−.203 (.216)	.031 (.118)
GDP^2	.053 (.368)	−.008 (.017)	−.008 (.018)	.229 (.317)	−.007 (.018)
R^2	.512	.548	.565	.629	.637
Obs	43	43	43	42	42

The dependent variable is the average Gini coefficient between 1980 and 2000. GDPand education are the initial values, stock market development is the average. All regressions include a dummy for Latin America. Coefficients are estimated with Ordinary Least Squares. Robust standard errors within parenthesis, 5% and 10% significant coefficients in bold and italics, respectively.

Table 4. Stock market development and income inequality

IV - cross-section - 1980-2000

	IV 1	IV 1	IV 1	IV 1	IV 2	IV 2	IV 2	IV 2
smcap	**.238** (.085)	1.246 (.873)			*.090* (.05)	.281 (.487)		
$smcap^2$		−1.158 (.993)				−.213 (.492)		
smpr			**.142** (.047)	.205 (.441)			**.097** (.048)	−.039 (.217)
$smpr^2$				−.057 (.401)				.109 (.186)
sec 25	**−.236** (.07)	**−.462** (.218)	**−.191** (.064)	**−.197** (.075)	**−.168** (.069)	−.221 (.152)	**−.167** (.066)	**−.148** (.070)
R^2	.429	-.228	.465	.505	.555	.509	.623	.639
Obs	69	69	68	68	43	43	42	42
$F-Test$ $(p-value)$	4.22 (.009)	4.22 (.009) 3.82 (.014)	5.91 (.001)	5.91 (.001) 3.41 (.023)	19.20 (.000)	14.48 (.000) 15.69 (.000)	11.44 (.000)	6.85 (.000) 6.25 (.000)
$Sargan$	.203	.751	.249	.084	.305	.485	.278	.411

The dependent variable is the Gini coefficient between 1980 and 2000, the regressors are initial GDP, GDP^2 and sec 25, and the period average stock market development. Coefficients are 2SLS estimates, stock market development instrumented with [uk, ge, fr legal origins] and [investor_pr, eff_jud, $(investor\_pr)^2$, $(eff\_jud)^2$] respectively in IV 1 and IV 2. Standard errors within parenthesis, 5% and 10% significant coefficients in bold and italics, respectively. P-values are reported for the first stage F-test and for the Sargan test of overidentifying restrictions. Latin America dummy included in all equations.

Table 5. Investor protection and stockmarket development

OLS - cross-section - 1980-2000

	smcap		*smpr*	
	OLS	OLS	OLS	OLS
sec 25	.197 (.253)	−.345 (.260)	.196 (.368)	−.222 (.325)
GDP	**1.328** (.503)	**1.831** (.596)	.115 (.744)	1.007 (.752)
GDP^2	*−1.310* (.741)	**−2.368** (.880)	−.896 (.1.105)	−1.739 (1.116)
Investor_pr		**.051** (.010)		**.052** (.012)
eff_jud		**.046** (.017)		*.036* (.020)
uk_lo	**.189** (.077)		**.339** (.112)	
fr_lo	.024 (.085)		.099 (.124)	
ge_lo	.099 (.099)		−.002 (.154)	
R^2	.419	.661	.244	.436
Obs	69	43	68	42

The dependent variable is stock market development between 1980 and 2000. Coefficient estimates from the first stage of columns 1, 5, 3, and 7 of Table 4. Standard errors within parenthesis, 5% and 10% significant coefficients in bold and italics, respectively.

Table 6. Stock market development and income inequality Sensitivity analysis

	1	2	3	4	5	6	7	8
	OLS	OLS	OLS	OLS	IV 1	IV 1	IV 1	IV 1
smcap	**.102** (.034)	.064 (.098)			**.229** (.085)	1.123 (.848)		
$smcap^2$		.043 (.110)				−.097 (.927)		
smpr			*.063* (.036)	**.171** (.061)			**.146** (.052)	.266 (.515)
$smpr^2$				−.084 (.055)				−.112 (.471)
R^2	.515	.516	.506	.533	.409	-.213	.401	.477
Obs	66	66	65	65	66	66	65	65
$F-test$ $(p-value)$					5.14 (.003)	5.14 (.003) 5.15 (.003)	6.09 (.001)	6.09 (.001) 3.29 (.027)
Sargan					.261	.693	.413	.167

The dependent variable is the latest available observation of Gini coefficient after 1985. GDP and education are initial values. Stock market development is 1980-2000 average. Cols 5-8 report 2SLS estimates with stock market development instrumented by [uk, fr, ge legal origins]. Robust standard errors within parenthesis, 5% and 10% significant coefficients in bold and italics, respectively. P-values are reported for the first stage F-test and for the Sargan test. Latin America dummy included in all equations.

Table 6b. Stock market development and income inequality Sensitivity analysis

	OLS	OLS	OLS	OLS
smcap	**.148** (.047)	−.033 (.125)		
smcap^2		.304 (.224)		
smpr			**.093** (.036)	.044 (.092)
smpr^2				.055 (.117)
R^2	.609	.625	.632	.635
Obs.	44	44	40	40

The dependent variable is the average Gini over 1985-2000. GDP and education are initial values. Stock market development is 1985 value Robust standard errors within parenthesis, 5% and 10% significant coefficients in bold and italics, respectively. Latin America dummy included in all equations.

Table 7. Stock market development and income inequality

Robustness analysis - cross-section - 1980-2000

	OLS	OLS	OLS	OLS	IV 1	IV 1	IV 1	IV 1
smcap	**.124** (.036)	**.227** (.097)			**.293** (.108)	1.273 (.989)		
smcap^2		−.116 (.111)				−1.111 (1.109)		
smpr			**.078** (.036)	**.195** (.055)			**.158** (.055)	.267 (.414)
smpr^2				*−.089* (.049)				−.098 (.369)
gov	−.026 (.082)	−.058 (.097)	−.076 (.073)	−.088 (.073)	−.061 (.151)	−.371 (.293)	−.151 (.112)	−.147 (.106)
open	−.029 (.033)	−.028 (.035)	−.005 (.027)	−.015 (.027)	−.085 (.072)	−.073 (.089)	−.023 (.033)	−.028 (.038)
R^2	.557	.564	.551	.584	.379	-.133	.452	.526
Obs	69	69	68	68	69	69	68	68
$F-test$ ($p-value$)					3.21 (.029)	3.21 (.029)	4.88 (.004)	4.88 (.004)
						2.95 (.040)		2.83 (.046)
Sargan					.286	.537	.271	.088

Dependent variable is the average Gini coefficient over 1980-2000. The other control variables are GDP, GDP^2 and *sec*25. Coefficients in cols IV 1 are 2SLS estimates with stock market development instrumented by [uk, fr, ge legal origins]. Robuststandard errors within parenthesis, 5% and 10% significant coefficients respectively in bold and italics. P-values are reported for the first stage F and the Sargan tests

Table 8. Stock market development and income inequality

static panel - 1976-2000

	FE	RE	FE	RE	FE	RE
smcap	**.147**	**.132**	**.111**	**.104**		
	(.036)	(.031)	(.041)	(.035)		
smcap^2	$-$**.041**	$-$**.036**	$-.028$	$-.025$		
	(.016)	(.015)	(.018)	(.016)		
smpr					*.029*	**.026**
					(.016)	(.005)
smpr^2					$-.001$	
					(.016)	
sec 25	$-$**.172**	$-$**.145**	$-$**.194**	$-$**.164**	$-$**.149**	$-$**.177**
	(.064)	(.048)	(.072)	(.052)	(.068)	(.049)
GDP	$-.168$	$-$**.163**	$-$*.179*	$-$*.147*	$-.078$	$-$*.129*
	(.105)	(.068)	(.106)	(.071)	(.119)	(.106)
GDP^2	*.102*	*.087*	*.109*	*.085*	.088	**.109**
	(.057)	(.045)	(.058)	(.047)	(.064)	(.048)
R^2	.227	.236	.239	.241	.243	.236
Observations	157	157	157	157	144	144
Hausman Test		.755		.807	.026	.425
Time FE	No	No	Yes	Yes	No	No
($F-test$)			(.162)	(.182)		

The dependent variable is the Gini coefficient. Sample of 52 (50) countries, non overlapping five-year observations spanning 1976-2000. GDP and sec25 are initial values, *smcap* and *smpr* are average ones. Standard errors in parenthesis, 5% and 10% significant coefficients in bold and italics, respectively. P-values of the Hausman tests are reported below RE estimates. P-values of the F-test for time fixed effects are reported in parenthesis.

Table 9. Stock market development and income inequality

Dynamic Panel Data - 1976-2000

	GMM	GMM	GMM	GMM	GMM	GMM	GMM	GMM
$d\log(smcap)$	**.079** (.029)	*.058* (.030)	.039 (.055)	−.062 (.105)				
$d[\log(smcap)]^2$			.048 (.055)	110 (.066)				
$d\log(smpr)$					**.062** (.022)	**.065** (.022)	**.202** (.066)	**.158** (.062)
$d[\log(smpr)]^2$							−**.121** (.048)	−*.075* (.045)
$d\log(\text{sec}\,25)$	−.186 (.139)	−.110 (.106)	−.167 (.136)	−.062 (.105)	−.137 (.104)	−.094 (.088)	−.138 (.93)	−.079 (.077)
$d\log(Gini_{t-1})$	**.519** (.141)	**.615** (.118)	**.518** (.139)	**.624** (.123)	**.596** (.111)	**.657** (.089)	**.569** (.115)	**.649** (.100)
$d\log(GDP)$	−.027 (.179)	.185 (.157)	−.065 (.179)	.087 (.184)	.178 (.189)	*.336* (.205)	.123 (.178)	.237 (.193)
$d[\log(GDP)]^2$	.025 (.225)	−.221 (.218)	.056 (.224)	−.137 (.213)	−.166 (.225)	−.347 (.225)	−.124 (.200)	−.259 (.212)
Sargan	.617	.645	.702	.974	551	.987	.739	.985
m_2	.940	.908	.774	.622	.344	.868	.363	.770
Observations	84	84	84	84	84	84	84	84
Time FE	No	Yes	No	Yes	No	Yes	No	Yes
$F-test$		(.068)		(.002)		(.153)		(.007)

Dependent variable is the log difference of Gini. Sample of 36 (32) countries, non-overlapping 5-year observations spanning 1976-2000. Estimation was performed with Arellano-Bover two-step system-GMM procedure. All regressors in difference are instrumented with their lagged levels, all levels with lagged differences. Coefficient estimates are from the first step. Standard errors are reported within parenthesis, 5% and 10% significant coefficients are respectively in bold and italics. P-values for F-test, Sargan and m_2 tests are from the second step.

Table 10. Stock market development and income inequality

static panel - 36 countries - 1976-2000

	FE	FE	FE	RE
smcap	**.103** (.045)	**.08** (.033)		
$smcap^2$	$-.021$ (.035)			
smpr			.043 (.041)	**.051** (.014)
$smpr^2$			.008 (.035)	
sec 25	$-$**.164** (.063)	$-$**.154** (.060)	$-$**.176** (.069)	$-$**.212** (.052)
GDP	$-.151$ (.120)	$-.139$ (.118)	$-.139$ (.132)	$-.045$ (.085)
GDP^2	.146 (.099)	.135 (.097)	.196 (.111)	.090 (.090)
R^2	.163	.159	.158	.139
Observations	125	125	112	112
Country FE	Yes	Yes	Yes	No
Hausman Test	.029	.014	.015	.127

The dependent variable is the Gini coefficient. Sample of 36 countries, non-overlapping five-year observations spanning 1976-2000. GDP and sec25 are initial values, smcap is the period average. Standard errors are reported within parenthesis, 5% and 10% significant coefficients are in bold and italics, respectively. P-values for the Hausman tests.

Table 11. Stock market development and income inequality - Robustness analysis

	FE	FE	FE	FE	GMM	GMM	GMM	GMM
smcap	**.124** (.047)	**.065** (.026)			.79 (.063)	**.077** (.031)		
$smcap^2$	**−.058** (.038)				.002 (.054)			
smpr			.061 (.039)	**.043** (.017)			**.172** (.062)	**.075** (.021)
$smpr^2$			−.018 (.035)				*−.085* (.045)	
gov	−.097 (.093)	.060 (.090)	.069 (.102)	.073 (.102)	.054 (.195)	.058 (.118)	.029 (.073)	.045 (.082)
trade	**.095** (.040)	**.076** (.039)	**.137** (.044)	**.748** (.286)	−.021 (.030)	−.019 (.029)	*−.037* (.019)	*−.041* (.022)
R^2	.224	.203	.260	.257				
$S\arg an$					.996	.999	.993	.943
m_2					.793	.842	.462	.454
Obs	125	125	112	112	84	84	84	84

Dependent variable is Gini in FE and RE columns, log difference of Gini in GMM. Samplesof non-overlapping 5-year observations spanning 1976-2000. The regressors of equations in FE (GMM) are the same as inTable 8 (9) plus (log difference of) government expenditure, trade and private credit as a ratio of GDP. FE are fixed and random effects regressions, chosen on the basis of specification tests, whose statistics are available upon request. GMM are Arellano-Bover two-step system-GMM estimations, where differences of all regressors are instrumented with lagged levels and levels with lagged differences. Coefficients are from the first step, p-values for Sargan and m_2 tests are from the second. Standard errors are reported within parenthesis, 5% and 10% significant coefficients in bold and italics.

Chapter 3

How Does Financial Liberalization Affect Economic Growth?*

1 Introduction

Academic economists and practitioners have long debated over the effects of financial globalization on growth. The removal of restrictions on international capital transactions has on some occasions been welcome as a growth opportunity and in others blamed for triggering financial instability and banking crises. Yet, this debate has not addressed the impact of financial liberalization on the sources of growth.[1] Does it affect investments in physical capital or total factor productivity (TFP), or both? If so, in which ways? This paper is a first attempt at answering these questions. Moreover, it helps understand whether financial globalization has growth or level effects and whether it brings convergence or divergence in growth rates across countries.

A wide literature has investigated the effects of international financial liberalization on GDP growth. The theoretical predictions are ambiguous. Some works suggest that, by promoting cross-country risk-diversification, financial liberalization fosters specialization, efficiency in capital allocation and growth (see, for instance, Acemoglu and Zilibotti, 1997 and Obstfeld, 1994). By generating international competition, it may also improve the functioning of domestic financial systems, with beneficial effects on savings and allocation (see Klein and Olivei, 1999 and Levine,

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[1] The only evidence in this direction is provided by Levine and Zervos (1998), who estimate the relation between the sources of growth and measures of stock market integration based on asset pricing models.

2001). On the other hand, financial liberalization may be harmful for growth in the presence of distortions. It may trigger financial instability, as well as misallocation of capital (see Eichengreen, 2001, for a survey), which are detrimental for macroeconomic performance. The empirical literature has not been able to resolve this theoretical controversy. Some studies (see, for instance, Grilli and Milesi-Ferretti, 1995, Kraay, 2000 and Rodrick, 1998) found that financial liberalization does not affect growth, others that the effect is positive (Levine, 2001, Bekaert et al., 2003 and Bonfiglioli and Mendicino, 2004), yet others that it is negative (Eichengreen and Leblang, 2003). Many authors show the effects to be heterogeneous across countries at different stages of institutional and economic development (see Bekaert et al, 2003, Chinn and Ito, 2003 and Edwards, 2001) and countries with different macroeconomic frameworks (Arteta Eichengreen and Wyplosz, 2001). Perhaps surprisingly, very little evidence exists on the effects of financial globalization on the various sources of growth.

In this paper, I separately address the effects of international financial liberalization on capital accumulation and TFP levels and growth rates. Financial liberalization, i.e. the removal of restrictions on international financial transactions, may affect productivity both directly and indirectly. As a direct effect, it is expected to generate international competition for funds, thereby driving capital towards the most productive projects. Indirectly, it may foster financial development which in turn positively affects productivity (see Beck et al., 2000).[2] The sign of the direct effect of financial liberalization on capital accumulation, through increased international competition, is ambiguous. For instance, Acemoglu et al. (2005) suggest that the effect of competition may vary depending on the distance of a country to the world technology frontier. Moreover, the overall effect of financial openness on the stock of capital may be ambiuguous, as capital reallocations may translate into net inflows for some countries and outflows for others.[3] Given the results in Beck et al. (2000), I expect the indirect effect through financial development to be weak.

As another indirect channel, however, financial liberalization may trigger financial instability and banking crises, as a wide literature points out (see Aizenmann,

[2]Financial development can be defined as the ability of a financial system to reduce information asymmetries between investors and borrowers, trade and diversify risk, mobilize and pool savings, and ease transactions. Removing restrictions on international financial transactions (financial liberalization) may affect the way a financial system carries over its functions, hence financial development.

[3]Alfaro et al. (2004) show that financial libralization does not significanly affect net capital flows, but did not examine the interaction between financial liberalization and productivity.

2001 for a survey on the evidence on financial liberalization and crises). Whatever the mechanism generating banking crises, such events may harm the ability of a financial system to provide the economy with credit. As a consequence, both investments in physical capital and innovation can be expected to slow down. In the worst scenario, even TFP might drop, due to the need for shutting down productive projects. I account for the effects of financial instability by controlling all regressions for banking crises. In this way, any indirect effect of liberalization through crises is removed from the estimates for the index of financial liberalization. I also estimate the joint effect of crises and liberalization to assess whether open capital account eases or worsens the recovery from bank crashes. Before going through these estimations, I explicitely address endogeneity between financial liberalization and banking crises by means of multinomial logit regressions.

I follow three methodologies to assess the effects of financial liberalization and banking crises on investments and productivity, and a fourth to address the link between liberalization and crises. I perform difference in difference estimation of the the impact of regime switches, between capital restrictions and openness, and between crises and normal times. I focus on investment and TFP levels, and I use a panel data with yearly observations from at most 93 countries over the period 1975-1999. Next, I estimate the same relationships using five-year averages. When studying the effects on TFP growth, I also investigate whether there is evidence of conditional convergence. I estimate an equation for TFP growth rates as a function of initial productivity and the other controls over a period of 25 year in a sample of 85 countries. To overcome problems of unobserved country-specific effects and endogeneity of regressors, I adopt the system GMM dynamic panel technique proposed by Arellano and Bover (1995) and Blundell and Bond (1998). To assess whether financial liberalization favors the occurrence of banking crises, I estimate logits and multinomial logits for an indicator distinguishing between systemic and borderline crises (see Caprio and Klingebiel, 2002). I use the annual 93-country-panel spanning between 1975 and 1999.

The main results are the following. (1) The effect of financial liberalization on TFP is positive and large in magnitude, while it is weak and non-robust on investments. (2) The impact on TFP is both on levels and and growth rates, implying that financial liberalization is able to spur GDP growth in the short as well as in the long run. (3) Financial liberalization raises only the probability of minor banking crises in developed countries. (4) Banking crises harm both capital accumulation and productivity. (5) Institutional and economic development amplify the positive effects of financial liberalization on productivity and limit the damages from banking crises. (6) Neither financial liberalization nor banking crises affect the speed of convergence in TFP growth rates.

The contribution of this paper is mainly related to three strands of literature. The literature on growth and development accounting has shown that a large share of cross-country differences in economic performance is driven by total factor productivity (TFP) rather than factor accumulation (physical and human capital).[4] Hall and Jones (1999) point out that a substantial share of GDP per worker variation is explained by differences in TFP and provide evidence that productivity is to a large extent determined by institutional factors. Klenow and Rodriguez-Clare (1997) show that also GDP growth differentials are mainly accounted for by differences in the growth rates of TFP. These results suggest that financial globalization may affects the wealth of nations through its impact on TFP, rather than factor accumulation, and that it may be important to distinsuish between the two channels.

Several authors suggest that financial development spurs GDP growth by fostering productivity growth, not only by raising the funds available for accumulation. Theoretical papers by Acemoglu, Aghion and Zilibotti (2005), Acemoglu and Zilibotti (1997), Aghion, Howitt and Mayer (2005b) among others show that financial development may relieve risky innovators from credit constraints, thereby fostering growth through technological change. While earlier contributions (e.g., Greenwood and Jovanovic, 1990) suggest that financial development fosters growth simply by increasing participation in production and risk pooling, in the later works the relationship is also driven by advances in productivity. King and Levine (1993), and, in more detail, Beck Levine and Loayza (2000) show evidence of a strong effect of financial development on TFP growth, and only a tenuous effect on physical capital accumulation.

My analysis of the joint effects of financial liberalization and banking crises on the sources of growth is also related to the literature on financial fragility and confronts with some of its predictions. For instance, Martin and Rey (2003) propose a model with multiple equilibria where financial liberalization raises asset prices, investments and income in emerging market, though leaving the poorest more prone to financial crises. In Tornell et al. (2004) banking crises may arise as a by-product of the higher growth generated by financial liberalization, in countries with credit market imperfections. Feijen and Perotti (2005) suggest that financial liberaliza-

[4] See Caselli (2005) for a survey on the develpment accounting literature, and Easterly and Levine (2001) for the stylized facts on development and growth accounting.

tion increases the likelyhood that the lobbying over the credit market accessibility generates financial fragility in equilibrium. Reinhart and Kaminsky (1999) provide evidence from a sample of 25 countries that financial liberalization has predictive power on banking crises. Kaminsky and Schmuckler (2002) show that this negative effect works dominates in the three-four years immediately after liberalization, then positive growth effects tend to emerge.

The remainder of the paper is organized as follows. Section 2 gives a brief overview on growth and development accounting, which leads on to the discussion of my empirical strategy. In section 3, I describe the dataset, with particular attention to the indicators of financial liberalization and banking crises, as well as the construction of the data for physical capital and TFP. Section 4 presents the econometric methodologies, and section 5 reports the results from the estimation of the equations for investments. Section 6 shows the evidence on level and growth rates of TFP and section 7 concludes.

2 The empirical strategy

The literature on growth and developing accounting takes as starting poing the Cobb Douglas specification for the aggregate production function,

$$Y = AK^{\alpha}\left(HL\right)^{1-\alpha}, \tag{3.1}$$

where K is the aggregate capital stock, L the number of workers and H their average human capital. The term A represents the efficiency in the use of factors, and corresponds to the notion of total factor productivity (TFP). Several contributions on development accounting (see Caselli, 2005 for a survey and Hall and Jones, 1999) have shown that a large share of the cross-country variation in GDP per worker, $\frac{Y}{L}$, is explained by differences in A. The works on growth accounting (see Easterly and Levine, 2001 and Klenow and Rodriguez-Clare, 1997), focusing on the following expression

$$\frac{\dot{Y}}{Y} = \frac{\dot{A}}{A} + \alpha\frac{\dot{K}}{K} + (1-\alpha)\left(\frac{\dot{H}}{H} + \frac{\dot{L}}{L}\right), \tag{3.2}$$

have shown that also cross-country differentials in GDP growth are to a large extent generated by differentials in productivity growth ($\frac{\dot{A}}{A}$).

All studies on the impact of financial liberalization and banking crises on growth have focused on $\frac{\dot{Y}}{Y}$, without assessing whether the effects are transmitted through factor accumulation or changes in productivity, or both. To grasp the relevance of the exercise proposed in this paper, consider the following growth regression:

$$dy_{it} = b_0 + b_1 y_{it-1} + b_2' \mathbf{Z}_{it} + b_3 FLIB_{it} + b_4 BC_{it} + u_{it}, \tag{3.3}$$

where $dy_{it} \equiv d\log(Y_{it})$ is the growth rate of GDP in country i, y_{it-1} is the logaritm of lagged GDP, $\mathbf{Z}_{it}$ is a vector of control variables, $FLIB_{it}$ and BC_{it} are indicators of financial liberalization and banking crises respectively, and u_{it} is the error term. Suppose the estimate for $\hat{b}_3$ is not significantly different from zero. This may reflect the absence of an effect of financial liberalization on any source of growth, as well as the presence of two countervailing effects on capital and TFP accumulation. Understanding what lies behind the effects on aggregate GDP growth may be crucial for policy purposes.

Various aspects of financial markets, such as volume, international liberalization and the occurrence of banking crises, may be expected to affect both physical capital accumulation and factor productivity. Beck et al. (2000) have shown evidence of a strong effect of financial depth on productivity, and a much weaker on capital accumulation.[5] Klein and Olivei (1999) and Levine (2001) find that financial liberalization fosters financial development. Should financial liberalization and banking crises affect investment and productivity only through the effect on the volume of credits, their impact on TFP and capital accumulation would thus be expected to be strong and weak respectively. However, there may be other, more direct effects as well.

Opening up the economy to capital inflows and outflows increases the degree of competition among international financial markets, which may lead to improvements in the allocative efficiency of the financial system. This implies that, holding financial depth constant, the average productivity of the financed projects might be higher than under autarky. Financial liberalization also allows for international risk-diversification, which may channel more resourses to risky innovation. Both effects may in turn shift resources away from physical capital accumulation towards TFP growth. As pointed out by Obstfeld (1994), financial globalization promotes specialization, just like trade, raising TFP where productivity is already high, and physical investments in countries far from the technology frontier.

Banking crises may hit industrial sectors to different extents. Financial insta-

[5] Financial depth is often used in the empirical literature as a measure of financial development, since it accounts for the weight of financial intermediation in the economy.

bility may induce the investors to take less risk, thereby shifting resources from innovation, which is typically riskier, to capital accumulation. However, the opposite might happen if a country deliberately invested in innovation to more quickly recover from the crisis.

3 The data

I perform the analysis on three datasets: a cross-section of 85 countries with data averaged over the period 1975 and 1999, and two unbalanced panels comprising up to 93 countries with annual and five-year observations over the period 1975-1999. As Table A shows, the largest sample includes twenty-two developed and seventy-one developing countries from all continents. The following subsections describe the main variables I include in the regressions.

3.1 Control variables

When assessing the effects of financial liberalization and banking crises on capital accumulation and productivity, I also control for a number of variables.

- Initial real per capita GDP (**rgdpch** from the PWT 6.1) accounts for different stages of economic development. It is often claimed that richer countries are more likely to have open financial markets, hence the effect of financial liberalization might seem spurious if initial GDP is not controlled for. If adding this variable to the regressions does not take away significance from the coefficient for financial liberalization, the suspects of spurousness are less sound.

- I include government expenditure as a ratio of GDP (**kg** from the PWT 6.1) in the regressions for capital accumulation. Several theories predict that government expenditure crowds out private investments. If this is the case, I should expect a negative coefficient in the equation for dlk.

- Financial depth, as proxied by the ratio of total credit to the private sector over GDP (**privo** from Beck and Demirguc-Kunt, 2001) and its growth rate give a measure of the external finance available to firms. Klein and Olivei (1999) and Levine (2001) show that financial liberalization promotes financial development, which may be expected to foster productivity more than capital accumulation, according to Beck et al. (2000). Bonfiglioli and Mendicino (2004) also find that banking crises have a negative effect on privo, mainly where institutions are weak. Controlling for financial depth in the equations for both investments and productivity helps disentangle the direct effects of liberalization and crises from the indirect ones through financial development.

A recent literature on financial fragility points out that crises may come along as by-products of sustained growth of the financial system (see Tornell et al. 2004). Feijen and Perotti (2005) suggest that equilibria with financial fragility and high participation in the financial market may arise where political accountability is not very high and wealth inequality is high. Including privo and its growth rate in the logit regressions for banking crises allows me to test a reduced form of these theoretical predictions.

- I also control for openness to trade, proxied by import plus export as a ratio of GDP (**openk** from the PWT 6.1). Trade may affect the efficency of an economy through several channels, such as specialization according to comparative advantage, access to larger markets with more product variety and increased competition. These effects may in turn stimulate both capital accumulation and productivity growth. However, the impact of trade may also depend on the distance of a country to the world technology frontier, as suggested by Acemoglu et al. (2005) and Aghion, Burgess, Redding and Zilibotti (2005a).

- Intellectual property right protection is expected to enhance productivity by giving incentives for innovation. This is controlled for by using the measure (**ipr**) by Ginarte and Park (1997), which is available for five-year periods from 1960 to 1990.

- Demirguc-Kunt and Detragiache (1997) show that the existence of explicit deposit insurance increases the likelihood of bank runs and thus crises of the banking sector. Hence, I include a measure of deposit insurance (**depins**) from Demirguc-Kunt and Sobaci (2000) in the logit analysis for banking crises.

- I also control for **inflation** (from the World Development Indicators) in the logit for banking crises. I take this variable as an indicator of bad macroeconomic policies, which are likely to make a country prone to crises.

- Finally, I use indicators of economic and institutional development to check for heterogeneity in the effects of financial liberalization and banking crises on both investments and productivity. In the cross-sectional estimates for TFP growth I explicitely control for institutional quality using the Government

Anti-Diversion Policy index (**gadp**, from Hall and Jones, 1999) as a proxy. As an indicator of economic development, I construct a dummy (**developing**) that takes value 1 if the country is defined as low or middle-low income in the World Development Indicators, and 0 otherwise. In the panel regressions, I use these indicators to split the sample and construct interactive terms.

3.2 Financial liberalization

I use two 0-1 indicators of financial liberalization, which rely on de iure criteria. The first one, CAL, is a dummy variable that takes value 0 if a country has held restrictions on capital account transactions during the year, and 1 otherwise. The existence of restrictions is classified on a 0-1 base by the IMF in its Annual Report on Exchange Arrangements and Exchange Restrictions (AREAER), which is available for a maximum of 212 countries over the period 1967- 1996. [6] This is the most commonly used indicator of international financial liberalization.

The second indicator relies on the chronology of official equity market liberalization, which is available in Bekaert et al. (2003) for 95 countries from 1980 onwards. It takes value 1 if international equity trading is allowed in a given country-year, and 0 otherwise. This dummy variable, EML, differs from CAL because it only accounts for equity market liberalization and not, for instance, credit market liberalization. As opposed to CAL, it does not allow for policy reversals: it labels a country as open ever since its first year of liberalization.

Factors affecting capital accumulation and productivity may also influence the decision of a country to liberalize financial markets. Moreover, there may be countries adopting such reforms either after reaching certain levels of investments and productivity, or with the purpose to attain them. This may raise concerns of omitted variables bias or even endogeneity, when estimating the effect of financial liberalization on capital accumulation and TFP. I tackle the issue by estimating the following logit on the annual panel dataset:

$$\Pr\left(FLIB\_r_{it} = 1\right) = \frac{e^{\beta_o + \beta_1 X_{it}}}{1 + e^{\beta_o + \beta_1 X_{it}}},$$

where $FLIB\_r_{it} \in \{CAL\_r,\, EML\_r\}$ is an indicator of the reforms observed in country i at time t, and $\mathbf{X}_{it}$ is a set of covariates. CAL_r equals 0 if there are no

[6] Classification methods have changed in 1996, so that there are now 13 separate indexes that can hardly be compared to the previous single indicator. Miniane (2000) harmonized the classifications, though for a limited number of countries, and over a short time span.

reforms, 1 if a switch into capital account liberalization occurs, -1 if the switch is out of it. EML_r does not admit reversals, thus it equals 1 in case of equity market liberalization reforms, and 0 otherwise. When the dependent variable is CAL_r, the estimation is performed with a multinomial logit.[7] All standard errors are robust and clustered by country. Following Bekaert et al. (2003), I include among the covariates a measure of institutional quality (gadp), lagged real GDP (rgdpch), government expenditure (kg), openness to trade (openk), financial depth (privo), inflation and GDP growth. I also control for economic development (developing) and continental dummies.

The results in Table B show the geographical component to capture reforms the most.[8] Perhaps surprisingly, the coefficient for gadp, not significantly different from zero, tells that financial liberalization is not more frequent in countries with good institutions than in the others.

3.3 Banking crises

Banking crises are subject to various classifications. I adopt a zero-one anecdotal indicator of bank crises, proposed by Caprio and Klingebiel (2003), who keep record of 117 systemic and 51 non-systemic crises occurring in 93 and 45 countries respectively, from the late 1970's and onwards. On a yearly base, the variable BC takes value 2 or 1 if the country has experienced a systemic or borderline banking crisis, respectively, and 0 otherwise. Caprio and Kilingebiel label a crisis as systemic if a great deal or all of a bank's capital has been exhausted and borderline if the losses were less severe. To make this definition criterion clearer, I refer to a few episodes. The 1991 crisis in Sweden as well as the 1998-99 crisis in Russia were systemic, since they involved insolvency or serious difficulties for 90 and 45 per cent of the banking system, respectively. The isolated failures of three UK banks between the eighties and the nineties, as well as the solvency problems of Credit Lyonnais in France in 1994-95, are instead labled as borderline crises.

Before going through the analysis of the effects of financial liberalization on the sources of growth, I address endogeneity between banking crises and financial

[7] All results are robust to the use of logit and probit on separate indicators: CAL_in (1 for switches into capital account liberalization, and 0 otherwise) and CAL_out (1 for switches out of capital account liberalization, and 0 otherwise).

[8] Note that, if I remove any of the continental dummies, the coefficients for the others remain significant.

liberalization, by estimating the following logit on the annual panel dataset:

$$\Pr\left(BC\_type_{it}=1\right)=\frac{e^{\beta_o+\beta_1\mathbf{X}_{it}+\gamma FLIB_{it}}}{1+e^{\beta_o+\beta_1\mathbf{X}_{it}+\gamma FLIB_{it}}}.$$

The variable BC_type_{it} takes value one if a banking crisis of a given *type* (systemic, borderline, or either one) has occurred in country i at time t. The vector $\mathbf{X}_{it}$ includes a series of covariates, and $FLIB_{it}$ is the binary indicator of financial liberalization. To appreciate the effects of all covariates, I also estimate a multinomial logit for BC_{it}, which takes values 1 and 2 in case of borderline and systemic crises respectively, and zero when no crises occur.[9] I cluster the standard errors by country.

Table C reports the results for BC_all, which equals 1 if any type of crisis has occurred, and 0 otherwise. Neither indicator of $FLIB$ has significant coefficient estimates. The variables raising the likelihood of crises the most are high inflation and the existence of explicit deposit insurance, as already shown by Demirguc-Kunt and Detragiache (1997). High real GDP per capita and growth rate of financial depth significantly reduce the probability of crisis. The first result is in line with the predictions in Martin and Rey (2004), while the second seems to contraddict the "bumpy path" hypothesis proposed by Tornell et al. (2004). Splitting the sample between developed and developing countries (columns 3-4 and 7-8), I find that CAL has a positive effect on the likelihood of banking crises in developed countries, while the growth rate of private credit is a more important factor in developing countries.

Finally, I exploit the classification in Caprio and Klingebiel (2002) and estimate with a multinomial logit the effects of all covariates on systemic versus borderline banking crises. Table D shows that CAL only has a positive effect on the likelihood of borderline banking crises in developed countries. The positive coefficient in column 3 of Table C is explained by the fact that most banking crises in developed countries are borderline. Deposit insurance, high real per capita GDP and the growth rate of financial depth mainly affect the probability of systemic crises. High inflation has opposite effects on the likelihood of the two types of crises: negative for borderline and positive for systemic crises. Equity market liberalization has no effect at all.

[9] I estimated the same model with pooled probit and fixed effects probit. Since the results are not sensitive to the estimation technique, I just report coefficients from the multinomial logit estimates.

3.4 Capital accumulation

I construct the series of the log-difference of physical capital stocks (dk) following the perpetual inventory method as in Hall and Jones (1999), using data from the Penn World Tables 6.1. I estimate the initial stock of capital, K_{t_0} as $\frac{I_{t_0}}{g+\delta}$, where g is the average geometric growth rate of total investments between t_0 and t_0+ 10.[10] In the paper t_0 is 1960, since I have data on investments dating back to that year for most countries.[11] A depreciation rate δ of 6 per cent in ten years is assumed. The later values of the capital stock are easily computed as $K_t = (1- \delta)K_{t-1}+ I_t$.

3.5 Productivity

I construct the series of total factor productivity following the Hall and Jones (1999) approach to the decomposition of output. I assume the production function in country i to be

$$Y_i = K_i^{\alpha} \left(A_i H_i L_i\right)^{1-\alpha},$$

where Y_i is the output produced in country i, K_i is the stock of physical capital in use, A_i is labor-augenting productivity, L_i is the labor in use (rgdpch* pop/ rgdpwok from the PWT 6.1), and H_i is a measure of the average human capital of workers ($H_i L_i$ is therefore human capital-augmented labor).[12] The factor share α is assumed constant across countries and equal to 1/3, which matches national account data for developed countries. I adopt the following specification for labor-augmenting human capital as a function of the years of schooling, s_i:

$$H_i = e^{\phi(s_i)}.$$

I rely on the results of Psacharopulos' (1994) survey and specify $\phi(s_i)$ as a piecewise linear function with coefficients 0.134 for the first four years of education, 0.101 for the next four years, and 0.068 for any value of $s_i > 8$.

Equipped with data on capital, output per worker, population and schooling

[10]Investments are defined as I = ki*rgdpch*pop from the PWT 6.1.

[11]In the countries which have no data for 1960 t_0 is the first year followed by at least 15 observations.

[12]In Hall and Jones (1999) Y_i is rgdpch*pop from the PWT, net of the value-added of the mining industry. Following Caselli (2005), I simplify and take rgdpch*pop.

(from Barro and Lee, 2001), I can compute the series of total factor productivity as

$$A_i = \frac{Y_i}{L_i}\frac{1}{H_i}\left(\frac{K_i}{Y_i}\right)^{-\frac{\alpha}{1-\alpha}}.$$

4 Econometric specifications and methodologies

In the next sections, I follow various methodologies to estimate the effects of financial liberalization and banking crises on the sources of growth. First, I fully exploit the cross-sectional and time-series information in the annual dataset and estimate

$$P_{it} = \beta_0 + \beta_1' \mathbf{X}_{it-1} + \gamma FLIB_{it-1} + \delta BC_{it-1} + \eta_i + \nu_t + \varepsilon_{it}, \qquad (3.4)$$

where P_{it} is a proxy for the outcome variable (either $\frac{\dot{K}}{K}$, $\frac{\dot{A}}{A}$ or $log(A)$ in the variuos specifications) observed in country i at year t, $\mathbf{X}$ are control variables, $FLIB$ is a dummy for financial liberalization and BC an indicator of banking crises. To reduce problems with simultaneity bias, all regressors enter as lagged values. η_i is a country-specific fixed effect capturing heterogeneity in the determinants of P that are specific to i. Its inclusion in (3.4) implies that γ is only estimated from the within-country variation around the liberalization date. The fixed year effects (ν_t) allow me to compare the change in P between the pre and post-reform periods in countries that have liberalized with the change in the countries that maintained the restrictions. This means that equation (3.4) is a "difference in difference" specification, since it implies differencing out the time-mean for each i, and the common trend for all i's at any t.

Two main problems may undermine the ability of γ to identify a causal link from financial liberalization to the sources of growth. First, there may be concerns about the selection of the countries that liberalized. As the results in Table B suggest, geographical location is a good predictor for reforms on international capital transactions. Suppose there are fewer liberalization episodes among countries of a certain area which also experiences particularly low productivity growth. This area-specific productivity trend may bias the effect of financial liberalization upwards. To control for this bias, I check if there are such differences across areas (Asia, Latin America, Africa, Europe+North America) and, if so, I include interacted time-area dummies. Table E reports the percentage of observations with capital account and equity market liberalization reforms (rows 1-2 and 4, respectively), the share of country-years with open capital and equity markets (rows 3 and 5), and the means of TPF (levels and growth) and capital accumulation across continents. Note from rows 1 and 2 that Africa, accounting for almost half of the sample, has the least number of capital account reforms and a very bad performance in terms of productivity growth. On the other hand, Europe and North America have the highest incidence of unreverted capital account liberalizations, the best performance in terms of productivity and the worst in capital accumulation. Moreover, in row 4, Asia has the highest number of equity market reforms and the highest average TFP growth. This suggest to control the difference in difference regressions for continental trends in both productivity and capital accumulation.

A problem of endogeneity of policy changes may also arise. Suppose a country opens up when experiencing an economic crisis to help the recovery or alternatively when it is already on a sustained growth path. This may attribute a negative or positive effect to financial liberalization which is actually due to a trend, thereby producing biased estimates. As a solution to this problem, I control for a dummy taking value 1 during the three or five years prior to the liberalization and zero otherwise. This allows me to verify whether the change in P was part of a previous trend or caused by liberalization.

To assess the effects of policy changes and banking crises in the medium-run, I also perform difference in difference estimates on a five-year panel dataset. In this case, the dependent variable is observed at the end of the period, while the regressors are expressed as beginning-of-period values.

When investingating TFP growth, I am also interested in the effects of liberalization along the transition. Therefore, I estimate the following productivity growth regression:

$$da_{i(t-\tau,t)} = \beta_0 + \lambda a_{it-\tau} + \beta_1' \mathbf{X}_{i(t-\tau,t)} + \gamma FLIB_{i(t-\tau,t)} + \delta BC_{i(t-\tau,t)} + u_{it}, \qquad (3.5)$$

where $da_{i(t-\tau,t)} = 100\frac{\log(A_{it})-\log(A_{it-\tau})}{\tau}$ and the regressors indexed by $(t-\tau,\, t)$ are τ-year period averages. A coefficient estimate $\hat{\lambda} < 0$ indicates that there is conditional convergence in productivity. The speed of convergence b can be obtained from the definition of $\lambda = -100\frac{1-e^{b\tau}}{\tau}$. I first estimate equation (3.5) on a 25-year cross section (τ= 25). As enphasized by the empirical growth literature (see Temple, 1999 for a survey), cross-sectional estimates have several limits. They do not allow me to exploit the time-series variation in the data, which is important to assess the effects of reforms, such as financial iberalization; nor to control for omitted variables, country-specific effects and endogeneity of the regressors. In this case, addressing endogeneity with an instrumental variable strategy looks rather difficult. Legal origins may be a good instrument for financial development (see La Porta et al, 1997), but do not look particularly suitable to instrument a variable as $FLIB$, which involves policy changes and perhaps reversals over the sample. Bekaert et al. (2003) address the issue by separately estimating a probit for $FLIB$, and find that the quality of institutions is crucial in determining the choice of liberalization. But as the institutional framework is known to be an important determinant of TFP (see, among others, Hall and Jones, 1999), it does not seem a valid instrument for $FLIB$, in a regression for TFP.

I address the first problem by turning to panel data. Note that the specification of equation (3.5) with $u_{it} = \eta_i + \nu_t + \varepsilon_{it}$ includes the lagged dependent variable. It follows that, even if ε_{it} is not correlated with $a_{it-\tau}$, the estimates are not consistent with a finite time span. Moreover, consistency may be undermined by the endogeneity of other explanatory variables, as in the cross-sectional estimates. To correct for the bias created by lagged endogenous variables, and the simultaneity of some regressors, I follow the approach proposed by Arellano and Bover (1995) and Blundell and Bond (1998). I estimate the following system with GMM

$$da_{it} = \beta_0 + \theta da_{it-5} + \boldsymbol{\beta}_1' d\mathbf{X}_{it} + \gamma dFLIB_{it} + \delta dBC_{it} + d\nu_t + d\varepsilon_{it} \quad (3.6)$$

$$a_{it} = \beta_0 + \theta a_{it-5} + \boldsymbol{\beta}_1' \mathbf{X}_{i(t-5,t)} + \gamma FLIB_{i(t-5,t)} + \delta BC_{i(t-5,t)} + \eta_i + \nu_t + \varepsilon_{it} \quad (3.7)$$

where da_{it} equals $\log(\frac{A_{it}}{A_{it-5}})$, and the other regressors are the same as in the previous equations. Levels indexed by $(t-5,\ t)$ are five-year averages. η_i, ν_t and ε_{it} are respectively the unobservable country- and time-specific effects, and the error term, respectively. The presence of country effect in equation (3.7) corrects the omitted variable bias. The differences in equation (3.6) and the instrumental variables estimation of the system are aimed at amending inconsistency problems. I instrument differences of the endogenous and predetermined variables with lagged levels in equation (3.6) and levels with differenced variables in equation (3.7). For instance, I take a_{it-15} as an instrument for da_{it-5} and $Flib_{it-10}$ for $dFLIB_{it}$ in (3.6) and da_{it-10} as an instrument for a_{it-5} and $dFLIB_{it-5}$ for $FLIB_{it}$ in (3.7). I estimate the system by two-step Generalized Method of Moments with moment conditions $E[da_{it-5s}\ (\varepsilon_{it} - \varepsilon_{it-5})] = 0$ for $s \geq 2$, and $E[dz_{it-5s}\ (\varepsilon_{it} - \varepsilon_{it-5})] = 0$ for $s \geq 2$ on the predetermined variables z, for equation (3.6); $E[da_{i,t-5s}\ (\eta_i + \varepsilon_{i,t})] = 0$ and $E[dz_{i,t-5s}$ $(\eta_i + \varepsilon_{i,t})] = 0$ for $s = 1$ for equation (3.7). I treat all regressors as predetermined. The validity of the instruments is guaranteed under the hypothesis that ε_{it} are not second order serially correlated. Coefficient estimates are consistent and efficient if both the moment conditions and the no-serial correlation are satisfied. To validate the estimated model, I apply a Sargan test of overidentifying restrictions, and a test of second-order serial correlation of the residuals. As pointed out by Arellano and Bond (1991), the estimates from the first step are more efficient, while the test statistics from the second step are more robust. Therefore, I will report coefficients and statistics from the first and second step respectively. Note that in this case the speed of convergence (divergence) is given by $\theta = e^{5b}$.

5 Financial liberalization, banking crises and capital accumulation

In this section, I estimate the following equation for investments

$$dk_{it} = \beta_0 + \beta_1' \mathbf{X}_{it-\tau} + \gamma FLIB_{it-\tau} + \delta BC_{it-\tau} + \eta_i + \nu_t + \varepsilon_{it},$$

where $dk_{it} = 100\frac{\log(K_{it})-\log(K_{it-\tau})}{\tau}$ proxies physical capital accumulation observed in country i at time t.[13] I take different frequencies, with τ equal to one and five years respectively, to assess the impact on the short and medium run. When I use the five-year panel, the dependent variable is observed at the end of the period and the regressors at the beginning. Since $FLIB$ is a binary indicator variable both in the annual and five-year panel, the coefficients will be difference in difference estimates.

Table 1a reports the results from the difference in difference regessions of dlk on yearly data. The specification in coulumn 1 only includes the indicators of capital account liberalization (CAL) and banking crises (BC), whose effects on investments are nil and negative, respectively. These coefficients are robust to controlling for trends in investments up to three years prior to capital account liberalization (CAL_switch3) and for time-continent effects, as reported in column 2.[14] Column 3 shows that banking crises have no different effect across financially open and restricted countries. When I control for real per capita GDP, government expenditure as a ratio of GDP and credit to the private sector as a ratio of GDP (column 4), CAL remains insignificant, while the negative coefficient for BC becomes only marginally

[13] The evidence is robust to the use of investments as a ratio of GDP as a proxy of the dependent variable. The results are availablie upon request.

[14] The results do not change if I use CAL_switch5, which equals 1 for the five years prior to the reform.

significant (it is different from zero at the ten per cent level). Note however that its significance is fully restored when any of the additional controls is removed from the regression (result not reported). The coefficients in column 4 show that richer countries accumulate more capital, while government expenditure tends to crowd out investments. The growth rate of physical capital is lower where financial intermediation (as proxied by *privo*) is higher and has grown less (the latter is not reported, but available upon request). This suggests that countries invest more in physical capital when their financial systems are at early stages of development and growing rapidly. Columns 5 and 6 report the estimates for the subsamples of developed and developing countries, as defined by the World Bank.[15] Interestingly, capital account liberalization has a positive effect on investments in the developed countries, and no impact in the others. As in column 4, removing any of the additional controls restores the negative coefficient for BC, without affecting the positive estimate for CAL in the developed countries. Finally, the results are robust to the inclusion of openness to trade, whose coefficient always turns out to be insignificant and is thus omitted.

In Table 1b I replicate the estimations of Table 1a replacing the capital account indicator with the indicator of equity market liberalization. All columns suggest that EML has a positive effect on capital accumulation, while the other regressors behave as in Table 1a.[16]

The difference in difference estimates from the five-year panel, reported in Tables 2a-2b, do not show any significant differences from the results obtained on the annual dataset. Capital account liberalization has almost no effect on investments, while equity market liberalization is generally investment-enhancing. Holding the other factors and TFP constant, these results would support the evidence in Bekaert et al (2003) that open equity markets promote GDP growth, while open capital account, as such, is not as effective.

[15]Heterogeneity in the effects of financial liberalization could also be addressed by including an interacted dummy $FLIB*$ *developing* in the full-sample regression. This method, however, may deliver biased estimates if there is heterogeneity in other coefficients, as shown in Tables 1a-1b.

[16]The estimation sample of Table 1b is a subset of the sample in Table1a. However, the coefficients for CAL are not sensitive to the sample. Results from re-estimating Table 1a on the sample of Table 1b are available upon request.

6 Financial liberalization, banking crises and productivity

In this section I estimate the effects of FLIB both on the level of TFP and its growth rate, which both contribute GDP growth. As pointed out by Klenow and Rodriguez-Clare (1997), any increase in productivity does not only raise output holding constant factor employment, but also fosters factor accumulation, which translates into higher GDP growth along the transition.

6.1 Level TFP: difference in difference estimates

I estimate the following equation for the logaritm of the level of TFP (a),

$$a_{it} = \beta_0 + \boldsymbol{\beta}_1' \mathbf{X}_{it-\tau} + \gamma FLIB_{it-\tau} + \delta BC_{it-\tau} + \eta_i + \nu_t + \varepsilon_{it},$$

in the panel datasets with annual and five-year data. When I use the five-year panel, the dependent variable is observed at the end of the period and the regressors at the beginning. As already mentioned in sections 4 and 5, this is a difference in difference specification.

Tables 3a and 3b report results from the yearly panel. The coefficients for CAL and EML are positive and significant across all specifications in columns 1-4. While equity market liberalization has a stronger effect in developing countries, the removal of capital account restrictions is beneficial in all countries, as shown by columns 5-6 of both tables. Banking crises have a negative and significant effect on TFP under all specifications. Note that when I add intellectual property rights protection among the regressors, twenty countries drop out of the sample due to missing observations. Nevertheless, the estimates for CAL, EML and BC in the equations of columns 1-3 do not change if I restrict the sample. Interestingly, the coefficients for $privo$ in columns 4-6 suggest that financial development on average tends to have a positive effect on productivity. However, its effect is positive in the developing countries and negative in the developed ones. This result may support the hypothesis that financial development favors convergence in productivity. Notice that the coefficients for financial liberalization and banking crises remain significant, even after controlling for financial development. This suggest that both have a direct effect on productivity. The coefficient estimates for ipr confirm the expectations of a positive effect on TFP, mainly in the developed countries where R&D capacity is probably higher.

In Tables 4a and 4b I report the results from the difference in difference estimates on the five-year panel. Here, the dependent variable is observed at the end of the five-year period, the dummy for financial liberalizaiton takes value 1 if a country has experienced no restrictions for at least one year and BC equals one if there has been at least one year of banking crisis. The positive coefficients for CAL is significant in the basic specification of column 1 and remains significant when I include pre-reform trends, continent-time effects and the full set of control variables. BC has a negative effect on TFP under every specification. The positive coefficient for equity market liberalization is more robust than that for CAL, and survives in most columns of Table 4b. Among the other control variables, the most significant is financial depth, which affects productivity positively in the developing countries, as in Tables 3a and 3b.

6.2 TFP growth and convergence

To evaluate the effects on productivity growth, I perform cross-sectional estimations of the following equation:

$$da_{i(t-25,t)} = \beta_0 + \lambda a_{it-25} + \beta_1' \mathbf{X}_{i(t-25,t)} + \gamma FLIB_{i(t-25,t)} + \delta BC_{i(t-25,t)} + \varepsilon_{it}.$$

The regressors indexed by $(t-25,\, t)$ are expressed in twenty-five-year averages. It follows that the estimates for γ and δ capture the effects of the occurrence and length of financial liberalization and banking crises on productivity growth. Period averages cannot, though, discriminate between liberalizations and crises happening early and late in the sample, nor between inerrupted and uninterrupted episodes amounting to the same mean.

The results in Tables 5a and 5b support the hypothesis of conditional convergence in productivity in robust way, with an implied speed of convergence b between 1 and 2 per cent per year.[17] The effect of banking crises on TFP growth is negative and significant under all specifications. In Table 5a, capital account liberalization has a positive and significant coefficient only under the basic specification (column1), and has no different effect across countries that experienced banking crises or and those that did not (column 2). The coefficient for $a_{t-25} * CAL$, aimed at assessing whether financial liberalization affects the pace of convergence, is nil in column 3. EML in Table 5b holds a positive and significant coefficient throughout columns

[17] Remember that the speed of convergence is computed from $\lambda = -100\frac{1-e^{25b}}{25}$.

1-3. Like CAL, it does not interact with banking crises nor with the initial level of productivity. It loses its significance once I control for GADP in columns 4 and 5. Both Table 5a and 5b suggest that the institutional factors captured by GADP, together with initial productivity, are the most important determinant of TFP growth. None of the other control variables seem to affect productivity growth.

The dynamic panel data estimates in Tables 6 and 7 confirm the cross sectional evidence in favor of conditional convergence in productivity. The implied speed of convergence is now higher and lies between 1.2 and 4.4 per cent per year. Both measures of $FLIB$ spur productivity growth in a robust way, while the negative effect of banking crises is now weaker. The coefficients for both CAL and EML lose significance only when I control for *privo* in columns 3 and 6. This suggests that the growth rate of TFP, as opposed to its level, is mostly affected by financial liberalization through financial development rather than directly. This evidence is consistent with the results obtained for GDP growth in Bonfiglioli and Mendicino (2004). Trade does not seem to have a significant effect on TFP growth.

Table 7 reports the results for the interactions of financial liberalization with banking crises, and the interaction of both $FLIB$ and BC with the level of economic development. Columns 1 and 2 show that banking crises and capital account liberalization do not affect the speed of convergence, while EML slows it down. Equity market liberalization has a larger benefit on the countries with higher initial productivity levels, which recalls the predictions in Aghion et al. (2005b) for financial development and Aghion et al. (2005a) for product market liberalization. The coefficients in columns 3 and 4 suggest that the joint effect of financial liberalization and banking crises harms productivity growth. Columns 5 and 6 show that BC lowers TFP growth everywhere, while $FLIB$ has positive effects in developed and negative effects in the developing countries. The same holds in columns 7 and 8, where I distinguish between countries with high and low institutional quality, as measured by GADP. These results support the existence of a robust positive effect of financial liberalization on productivity. Arguably, the threat of an increase in competition for funds from abroad favors the channeling of resources towards innovative projects raising aggregate TFP.

7 Conclusions

A wide literature has focused on the effect of financial liberalization on GDP growth, often finding mixed results. To better understand the effect of financial liberalization, however, it is important to know the channels through which it operates. This paper has attempted to probe deeper into the relationship by separately studying the impact of financial openness on two sources of income growth: capital accumulation and productivity. Contrary to the existing literature, I find fairly robust results. In particular, financial liberalization has little effect on capital accumulation, while it has a strong positive effect on productivity. Financial liberalization appears to spur TFP growth through financial development, while it has a direct impact on the productivity level.

The paper has also studied the impact of financial instability on economic performance and the relationship between financial openness and crisis. As expected, crises are found to be detrimental, both for productivity and capital accumulation. However, there is no evidence that financial openness increases the likelihood of crisis, except for borderline crisis in developing countries. Thus, the concern that the removal of barriers to capital mobility may expose an economy to higher financial risk seems unwarranted.

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Table A

Countries, samples and financial liberalization dates

Country	Panel	Cross	CAL_on	CAL_off	EML_on	Country	Panel	Cross	CAL_on	CAL_off	EML_on
Algeria	x					Denmark	x	x	1988		
Angola	x					Ecuador	x	x	1971-1988-1995	1970-1986-1993	1994
Argentina	x	x	1967-1993	1970	1989	Egypt	x	x			1992
Australia	x	x	1984			El Salvador	x	x			
Austria	x	x	1991			Equatorial Guinea	x				
Bangladesh	x	x			1991	Ethiopia	x	x			
Benin	x	x	1967	1968		Finland	x	x	1991		
Bolivia	x	x	1986	1981		France	x	x	1990	1968	
Botswana	x	x			1990	Gabon	x	x	1967	1968	
Brasil	x	x			1991	Gambia	x	x	1991		
Burkina Faso	x	x	1967	1969		Germany	x				
Burundi	x	x				Ghana	x	x			1993
Cameroon	x	x	1967	1968		Greece	x	x			1987
Canada	x	x				Guatemala	x	x	1973-1989	1980	
Cape Verde	x	x				Guinea	x	x			
Central Africa	x	x	1967	1968		Guinea Bissau	x	x			
Chad	x	x	1967	1968		Hong Kong	x	x			
Chile	x	x			1992	Iceland	x	x			1991
Colombia	x	x			1991	India	x	x			1992
Congo	x	x	1967	1968		Indonesia	x	x	1969		1989
Costa Rica	x	x	1980-1995	1974-1982		Israel	x	x			1993
Cote d'Ivoire	x	x	1967	1968	1995	Italy	x	x	1990		

Note. CAL_on and CAL_off report the dates of removal and adoption, respectively, of restrictions on capital account transactions. EML_on reports the dates of official liberalization of the equity market.

Table A

Countries, samples and financial liberalization dates

Country	Panel	Cross	CAL_on	CAL_off	EML_on	Country	Panel	Cross	CAL_on	CAL_off	EML_on
Jamaica	x	x			1991	Philippines	x	x		1969	1991
Japan	x	x	1979	1995	1983	Portugal	x	x	1993		1986
Jordan	x	x			1995	Rwanda	x	x			
Kenia	x	x			1995	Senegal	x	x	1967	1968	
Korea	x	x			1992	Sierra Leone	x				
Lesotho	x	x				Singapore	x		1978		
Madagascar	x	x	1967	1968		South Africa	x	x			1996
Malaysia	x	x	1973		1988	Spain	x	x	1994		1985
Mali	x	x				Sri Lanka	x	x			1991
Mauritania	x	x	1967	1968		Sweden	x	x	1993		
Mauritius	x				1994	Switzerland	x	x			
Mexico	x	x		1982	1989	Tanzania	x	x			
Morocco	x	x			1988	Thailand	x	x			1987
Mozambique	x	x				Togo	x	x	1967	1968	
Nepal	x	x				Trinidad and Tobago	x	x	1994		1997
Netherlands	x	x				Tunisia	x	x			1995
New Zealand	x	x	1984		1987	Turkey	x	x			1989
Nicaragua	x	x		1978		Uganda	x	x			
Niger	x	x	1967-1995	1968		United Kingdom	x	x	1979		
Nigeria	x	x			1995	United States	x	x			
Norway	x	x	1995			Uruguay	x	x	1978	1968-1993	
Panama	x	x				Venezuela	x	x		1984	1990
Papua New Guinea	x	x				Zambia	x	x			
Paraguay	x	x	1982	1984		Zimbabwe	x	x			1993
Peru	x	x	1978-1993	1970-1984	1992						

Note. CAL_on and CAL_off report the dates of removal and adoption, respectively, of restrictions on capital account transactions. EML_on reports the dates of official liberalization of the equity market.

Table B
Financial liberalization - yearly panel - logit and multinomial logit

	CAL_in	CAL_out	EML_in
Africa	15.426 ***	-15.299 **	-0.759 **
	4.508	7.287	0.318
Asia	15.653 ***	19.469 ***	-0.713
	4.534	7.177	0.452
Latin America	17.326 ***	22.334 ***	-0.980 ***
	4.585	7.308	0.344
Europe & N. America	15.592 ***	-17.587 **	-3.379 ***
	4.644	7.884	1.073
developing	0.304	-0.198	0.072
	0.394	1.002	0.291
gadp	3.333	0.317	1.226
	2.223	3.419	1.219
growth	1.041	-7.302	2.546
	5.418	4.758	3.582
inflation	-0.013 *	-0.004	0.000
	0.007	0.003	0.000
kg	0.148	-0.576	-0.265
	0.397	0.481	0.310
openk	0.237	0.721	0.149
	0.277	0.532	0.239
privo	-0.533 **	0.085	-0.049
	0.261	0.507	0.235
rgdpch	0.518	1.051	0.077
	0.601	0.936	0.260

Note. CAL_in and CAL_out indicate switches on and off capital account liberalization, respectively. The coefficients in theese columns are estimated with multinomial logit. EML_in indicates reforms of equity market liberalization. The coefficients in this column are estimated with logit. Africa, Asia, Latin America and Europe & N. America are continental dummies. Developing is a dummy for developing countries as defined by the World Bank. The variables growth, inflation, gov, open, privo and rgdp enter as lagged values. A constant is included in all regressions. The robust standard errors are clustered by country. *, ** and *** indicate that a coefficiant is significant at 10, 5 and 1 per cent level, respectively.

Table C

Financial liberalization and banking crises - yearly panel - logit

	1	2	3	4	5	6	7	8
			Developed	Developing			Developed	Developing
CAL	0.213	0.202	1.587 ***	-0.125				
	0.269	0.266	0.365	0.308				
EML					-0.030	0.047	0.465	-0.012
					0.339	0.322	0.799	0.362
depins	0.640 **	0.577 *	0.732	0.807 **	0.419	0.413	0.666	0.513
	0.307	0.305	0.523	0.415	0.295	0.301	0.435	0.399
rgdpch	-0.410	-0.659 **	-0.907 *	-0.941 **	-0.913 ***	-0.853 ***	-0.626 *	-1.063 **
	0.304	0.283	0.524	0.451	0.331	0.295	0.334	0.487
inflation	0.000 *	0.001 *	0.007 ***	0.000 **	0.000 *	0.000 *	0.006 ***	0.000
	0.000	0.000	0.003	0.000	0.000	0.000	0.002	0.000
openk	0.001	0.002	0.016 *	0.000	0.001	0.001	0.003	0.001
	0.003	0.003	0.009	0.004	0.004	0.004	0.010	0.004
privo	-0.737				0.111			
	0.699				0.693			
dprivo		-1.817 ***	0.970	-2.926 ***		-1.237 **	0.809	-2.033 ***
		0.515	1.647	0.717		0.505	1.762	0.767
Obs	1117	283	961	830	961	952	240	712

Note. The dependent variable is a binary indicator of banking crises (*BC_all*), that equals 1 if a crisis occurs and 0 otherwise. All regressors are in lagged values. Standard errors are clustered by country. *, ** and *** indicate that a coefficient is significant at 10, 5 and 1 per cent, respectively.

Table D

Financial liberalization and banking crises - yearly panel - mlogit

	1	2	3	4	5	6	7	8	9	10	11	12
			Developed		Developing				Developed		Developing	
	BL	SYS	BL	SYS	BL	SYS	BL	SYS	BL	SYS	BL	SYS
CAL	0.716 *	-0.187	1.921 **	0.641	0.335	-0.330						
	0.406	0.366	0.804	0.791	0.577	0.399						
EML							0.429	-0.159	0.579	0.406	0.942	-0.258
							0.702	0.405	1.121	1.086	0.748	0.437
depins	-0.200	0.767 *	2.397 ***	-0.844	-1.151	1.216 ***	-0.061	0.603	2.015 ***	-0.204	-1.307	0.872 **
	0.495	0.428	0.674	1.630	1.025	0.447	0.507	0.402	0.595	0.757	1.048	0.419
rrgdpch	0.204	-1.168 **	0.033	-2.442 **	0.013	-1.143 **	-0.204	-1.241 ***	-0.015	-2.565 ***	-0.894	-1.092 **
	0.410	0.496	0.446	1.029	0.752	0.575	0.623	0.443	0.515	0.648	1.145	0.512
inflation	-0.025 *	0.001 **	-0.015	0.009 ***	-0.018	0.000 **	-0.015 *	0.001 *	-0.004	0.010 ***	-0.010	0.000
	0.013	0.000	0.029	0.003	0.012	0.000	0.008	0.000	0.011	0.004	0.008	0.000
openk	0.007	0.000	0.040 ***	0.014	0.008	-0.001	0.005	0.000	0.023 *	-0.015	0.010	0.000
	0.005	0.005	0.013	0.016	0.006	0.005	0.006	0.005	0.012	0.012	0.008	0.005
grprivo	-1.423	-2.013 ***	3.642 **	0.925	-2.118 **	-3.261 ***	-0.133	-1.510 ***	2.556	0.641	-0.869	-2.294 ***
	1.126	0.538	1.769	2.166	1.074	0.788	1.117	0.512	2.186	1.850	1.348	0.791

Note. The dependent variable is an indicator of banking crises (*BC*), that equals 2 if a systemic crisis (SYS) occurs, 1 if the crisis is borderline (BL), and 0 otherwise. All regressors are in lagged values. The estimation is performed with multinomial logit. Standard errors are clustered by country. *, ** and *** indicate that a coefficient is significant at 10, 5 and 1 per cent, respectively.

Table E
Reforms and financial liberalization across continents

	Asia	Africa	Latin America	Europe & N. America
CAL_in	1.37	0.43	3.22	2.95
CAL_out	0.34	0	2.89	0
CAL	41.16	1.29	28.94	43.51
EML_in	5	3.41	3.81	1.14
EML	50	53	17.41	74.19
Level TFP	1.116	1.547	1.864	2.084
TFP growth	-0.114	-2.286	-2.559	-0.207
Capital accumulation	6.884	4.223	3.182	3.167
Observations	294	699	311	239

Note. The table reports the share (%) of observations with capital account and equity market liberalization (CAL and EML, respectively), switches into and out of capital account liberalization (CAL_in and CAL_out), and into equity market liberalization (EML_in). For the other variables, means are reported.

Table 1a

Capital account liberalization and capital accumulation - yearly panel - difference in difference

	1	2	3	4	5 Developed	6 Developing
CAL	0.700	0.412	0.273	0.528	1.956 **	0.099
	0.623	0.761	0.802	0.955	0.779	1.340
BC	-0.782 ***	-0.702 ***	-0.754 ***	-0.500 *	-0.473	-0.496
	0.217	0.224	0.243	0.305	0.304	0.403
CAL_BC			0.326			
			0.599			
lkg				-2.528 ***	-1.673 **	-3.700 ***
				0.840	0.744	1.227
lprivo				-1.021 *	-1.239 **	-1.343 *
				0.610	0.566	0.831
lrgdpch				5.036 ***	2.573	5.426 **
				1.668	1.691	2.189
CAL_switch3		-0.319	-0.314	0.043	-0.060	-0.113
		0.707	0.707	0.891	0.874	1.174
Time-continent	No	Yes	Yes	Yes	Yes	Yes
Obs	1900	1900	1900	1385	361	1024
Countries	93	93	93	79	20	59

Table 1b

Equity market liberalization and capital accumulation - yearly panel - difference in difference

	1	2	3	4	5 Developed	6 Developing
EML	0.995 ***	0.965 ***	1.066 ***	0.629 *	1.446 **	0.687 *
	0.242	0.315	0.336	0.339	0.631	0.420
BC	-0.664 ***	-0.483 ***	-0.436 ***	-0.341 ***	-0.263	-0.150
	0.092	0.100	0.114	0.107	0.220	0.131
EML_BC			-0.204			
			0.237			
lkg				-1.007 ***	-2.517 ***	0.191
				0.338	0.520	0.485
lprivo				-0.501 **	-0.223	-1.192 ***
				0.254	0.432	0.345
lrgdpch				3.511 ***	3.890 **	3.858 ***
				0.777	1.523	0.963
EML_switch3		0.449	0.457	0.422	1.209 **	0.328
		0.284	0.284	0.298	0.548	0.361
Time-continent	No	Yes	Yes	Yes	Yes	Yes
Obs	1482	1248	1248	1026	286	740
Countries	78	78	78	69	18	51

Note. The dependent variable is the annual growth rate of physical capital stock (dk). All regressors are in lagged values. The variables CAL_switch3 and EML_switch3 equal 1 in the 3 years prior to capital account and equity market reforms, respectively. The sample spans between 1975 and 1999. All regressions include a constant. Standard errors are clustered by country. *, ** and *** indicate that a coefficient is significant at 10, 5 and 1 per cent, respectively.

Table 2a

Capital account liberalization and capital accumulation - yearly panel - difference in difference

	1	2	3	4	5	6
					Developed	Developing
CAL	0.266	0.425	-0.438	0.658	0.703	0.846
	0.497	0.559	0.637	0.503	0.447	0.755
BC	-0.005 ***	-0.899 ***	-0.207 ***	-0.640 **	-0.381	-0.608 *
	0.281	0.289	0.340	0.266	0.266	0.364
CAL_BC			0.383 *			
			0.608			
lkg				-0.684	-0.054 **	-0.364
				0.496	0.503	0.662
lprivo				0.793	0.608	0.050
				0.377	0.758	0.449
lrgdpch				-0.589	-2.697 *	-0.676
				0.850	0.584	0.046
CAL_switch5		-0.540	-0.550	-0.744 *	-0.469	-0.033
		0.477	0.475	0.425	0.398	0.630
Time-continent	No	Yes	Yes	Yes	Yes	Yes
Obs	457	457	457	353	98	255
Countries	93	93	93	85	22	63

Table 2b

Equity market liberalization and capital accumulation - yearly panel - difference in difference

	1	2	3	4	5	6
					Developed	Developing
EML	0.604	0.786 *	0.583	0.830	-0.168	0.401
	0.395	0.470	0.581	0.591	0.877	0.801
BC	-0.722 ***	-0.589 **	-0.724 **	-0.471 *	-0.149	-0.550
	0.249	0.249	0.337	0.264	0.283	0.364
EML_BC			0.318			
			0.535			
lkg				-1.271 **	-0.910	-1.214
				0.548	0.558	0.790
lprivo				-0.150	1.360	-0.157
				0.449	0.956	0.551
lrgdpch				-0.759	-2.199	-0.867
				1.118	2.197	1.413
EML_switch5		-1.345	-1.262	-2.392	-1.626	0.281
		2.295	2.303	2.661	3.327	3.689
Time-continent	No	Yes	Yes	Yes	Yes	Yes
Obs	312	312	312	268	80	188
Countries	78	78	78	73	21	52

Note. The dependent variable is the 5-year average annual growth rate of physical capital stock (dk). All control variables are observed at the beginning of the period. CAL and EML equal 1 if liberalization is observed for at least one year in the period. The variables CAL_switch5 and EML_switch5 equal 1 in the 5-year period prior to capital account and equity market reforms, respectively. The sample spans between 1975 and 1999. All regressions include a constant. Standard errors are clustereded by country. *, ** and *** indicate that a coefficient is significant at 10, 5 and 1 per cent, respectively.

Table 3a

Capital account liberalization and level TFP - yearly panel - difference in difference

	1	2	3	4	5 Developed	6 Developing
CAL	0.140 ***	0.054 **	0.048 **	0.104 ***	0.104 **	0.123 ***
	0.020	0.022	0.023	0.023	0.047	0.030
BC	-0.063 ***	-0.053 ***	-0.055 ***	-0.057 ***	-0.108 ***	-0.049 ***
	0.007	0.007	0.007	0.007	0.018	0.008
CAL_BC			0.016			
			0.017			
lprivo				0.031 **	-0.069 **	0.068 ***
				0.015	0.032	0.018
lopenk				-0.013	0.078	-0.023
				0.022	0.102	0.023
ipr				0.016 *	0.042 **	0.005
				0.009	0.019	0.011
CAL_switch3		-0.003	-0.003	0.026	-0.022	0.040 *
		0.020	0.020	0.019	0.047	0.022
Time-continent	No	Yes	Yes	Yes	Yes	Yes
Obs	1844	1844	1844	1119	309	810
Countries	93	93	93	73	18	55

Table 3b

Equity market liberalization and level TFP - yearly panel - difference in difference

	1	2	3	4	5 Developed	6 Developing
EML	0.112 ***	0.111 ***	0.096 ***	0.071 ***	0.015	0.080 ***
	0.016	0.019	0.020	0.023	0.060	0.026
BC	-0.047 ***	-0.042 ***	-0.050 ***	-0.050 ***	-0.091 ***	-0.041 ***
	0.006	0.006	0.007	0.007	0.019	0.008
EML_BC			0.031 *			
			0.014			
lprivo				0.009	-0.061 *	0.046 **
				0.017	0.034	0.023
lopenk				-0.008	0.028	-0.015
				0.027	0.134	0.028
ipr				0.014	0.063 ***	-0.009
				0.009	0.019	0.011
EML_switch3		0.024	0.023	0.001	-0.025	0.011
		0.017	0.017	0.017	0.043	0.020
Time-continent	No	Yes	Yes	Yes	Yes	Yes
Obs	1451	1224	1224	814	239	575
Countries	78	78	78	67	18	49

Note. The dependent variable is the logaritm of TFP level (a). All regressors are in lagged values. The variables CAL_switch3 and EML_switch3 equal 1 in the 3 years prior to capital account and equity market reforms, respectively. The sample spans between 1975 and 1999. All regressions include a constant. Standard errors are clustered by country. *, ** and *** indicate that a coefficient is significant at 10, 5 and 1 per cent, respectively.

Table 4

Capital account liberalization and level TFP - 5-year panel - difference in difference

	1	2	3	4	5	6
					Developed	Developing
CAL	0.132 ***	0.070	0.073	0.121 **	0.032	0.063
	0.047	0.048	0.056	0.058	0.064	0.085
BC	-0.093 ***	-0.075 ***	-0.073 **	-0.112 ***	-0.032	-0.112 ***
	0.027	0.026	0.030	0.029	0.030	0.039
CAL_BC			-0.007			
			0.054			
lprivo				0.080 **	-0.069	0.103 **
				0.035	0.059	0.041
lopenk				0.002	-0.011	0.046
				0.055	0.042	0.082
ipr				-0.013	0.193 **	-0.044
				0.089	0.085	0.119
CAL_switch5		-0.046	-0.046	-0.036	-0.030	-0.045
		0.041	0.041	0.041	0.041	0.061
Time-continent	No	Yes	Yes	Yes	Yes	Yes
Obs	443	443	443	238	71	167
Countries	93	93	93	78	20	48

Table 4b

Equity market liberalization and level TFP - 5-year panel - difference in difference

	1	2	3	4	5	6
					Developed	Developing
EML	0.086 **	0.120 ***	0.070	0.175 ***	-0.013	0.094
	0.040	0.045	0.054	0.066	0.096	0.090
BC	-0.091 ***	-0.073 ***	-0.107 ***	-0.099 ***	-0.015	-0.123 ***
	0.025	0.023	0.032	0.031	0.028	0.043
EML_BC			0.080			
			0.050			
lprivo				0.121 ***	0.067	0.149 ***
				0.046	0.083	0.055
lopenk				0.004	0.006	0.039
				0.067	0.053	0.100
ipr				0.078	0.494 ***	-0.032
				0.139	0.158	0.173
EML_switch5		-0.289	-0.270	-0.573 *	-0.062	-0.232
		0.216	0.215	0.304	0.358	0.416
Time-continent	No	Yes	Yes	Yes	Yes	Yes
Obs	304	304	304	178	56	122
Countries	78	78	78	64	20	44

Note. The dependent variable is the 5-year average logaritm of TFP level (a). All control variables are observed at the beginning of the period. CAL and EML equal 1 if liberalization is observed for at least one year in the period. The variables CAL_switch5 and EML_switch5 equal 1 in the 5-year period prior to capital account and equity market reforms, respectively. The sample spans between 1975 and 1999. All regressions include a constant. Standard errors are clustereded by country. *, ** and *** indicate that a coefficient is significant at 10, 5 and 1 per cent, respectively.

Table 5a
Capital account liberalization and TFP Growth - cross-section

	1	2	3	4	5
a_25	-1.109 **	-1.102 **	-1.097 **	-1.374 ***	-1.563 ***
	0.448	0.446	0.480	0.433	0.416
CAL	1.326 *	0.577	1.523	0.236	-0.214
	0.713	1.059	1.876	0.590	0.654
BC	-4.134 ***	-4.464 ***	-4.144 ***	-3.587 ***	-3.961 ***
	1.384	1.517	1.387	1.202	1.277
CAL_BC		2.098			
		2.712			
a_CAL			-0.112		
			0.932		
gadp				7.074 ***	7.697 ***
				1.372	1.884
lprivo					0.302
					0.514
lopenk					-0.073
					0.508
ipr					-0.430
					0.576
R2	0.186	0.190	0.186	0.340	0.485
Obs	85	85	85	85	73

Table 5b
Equity market liberalization and TFP Growth - cross-section

	1	2	3	4	5
a_25	-0.871 *	-0.868 *	-0.928 *	-1.299 ***	-1.215 ***
	0.463	0.474	0.520	0.411	0.439
EML	2.380 ***	2.797 ***	1.869	0.117	-0.040
	0.641	1.016	1.464	0.710	0.665
BC	-2.501 *	-2.063	-2.435 *	-2.448 **	-2.801 **
	1.353	1.898	1.356	1.125	1.337
EML_BC		-1.073			
		2.118			
a_EML			0.298		
			0.829		
gadp				8.320 ***	8.361 ***
				1.603	2.351
lprivo					0.083
					0.625
lopenk					-0.064
					0.492
ipr					-0.508
					0.550
R2	0.236	0.238	0.237	0.409	0.432
Obs	72	72	72	72	65

Note. The dependent variable is the 25-year average annual growth rate of TFP (*da*). All regressors are expressed as period average, except for the logaritm of the initial TFP level. The sample spans between 1975 and 1999. All regressions include a constant. Robust standard errors are reported below the coefficients. *, ** and *** indicate that a coefficient is significant at 10, 5 and 1 per cent, respectively.

Table 6

TFP Growth - Dynamic Pane Data - System GMM

	1	2	3	4	5	6
da_1	0.834 ***	0.899 ***	0.893 ***	0.911 ***	0.890 ***	0.936 ***
	0.089	0.069	0.050	0.083	0.072	0.038
dCAL	0.133 ***	0.136 ***	0.073			
	0.050	0.052	0.053			
dEML				0.027	0.021	-0.038
				0.054	0.057	0.072
dBC	-0.064	-0.048	-0.079 **	-0.035	-0.075 **	-0.082 **
	0.040	0.035	0.036	0.045	0.036	0.039
dlopenk		-0.038	0.031		0.051	-0.048
		0.084	0.085		0.112	0.056
dlprivo			0.068 **			0.046
			0.028			0.032
Sargan (pvalue)	0.670	0.727	0.472	0.352	0.642	0.559
m2 (pvalue)	0.843	0.757	0.487	0.490	0.822	0.885
Time FE	Yes	Yes	Yes	Yes	Yes	Yes
Obs	433	371	329	301	263	253
Countries	89	78	75	76	67	67

Note. The dependent variable is the 5-year log-difference of TFP level (*da*). All regressors are 5-year period averages. The sample spans between 1975 and 1999. All regressions include a constant. The estimation is performed with the two-step system-GMM procedure. Coefficients and standard errors are reported from the first step. *, ** and *** indicate that a coefficient is significant at 10, 5 and 1 per cent, respectively. The p-values for the Sargan overidentification test and the second order serial correlation (m2) test are reported from the second step.

Table 7

TFP Growth - Dynamic Panel Data - System GMM

	1	2	3	4	5	6	7	8
da_1	0.917 ***	0.866 ***	0.853 ***	0.879 ***	0.800 ***	0.853 ***	0.835 ***	0.858 ***
	0.054	0.063	0.090	0.082	0.096	0.093	0.091	0.096
dCAL	-0.078		0.155 ***		0.310 ***		0.405 ***	
	0.133		0.049		0.098		0.122	
dEML		-0.153 *		0.249 *		0.192		0.378 **
		0.088		0.104		0.138		0.172
dBC	-0.051	-0.095	-0.044	0.036	-0.071	-0.040	0.029	0.118
	0.061	0.067	0.039	0.071	0.046	0.062	0.188	0.263
da_CAL	0.073							
	0.068							
da_EML		0.132 **						
		0.052						
da_BC	0.003	0.026						
	0.038	0.043						
dCAL_BC			-0.197 **					
			0.080					
dEML_BC				-0.326 ***				
				0.125				
dCAL_dev'ing					-0.491 **			
					0.208			
dEML_dev'ing						-0.239 *		
						0.140		
dBC_dev'ed					0.041	0.101		
					0.119	0.133		
dCAL_(1-gadp)							-1.220 **	
							0.498	
dEML_(1-gadp)								-0.914 **
								0.361
dBC_gadp							-0.163	-0.226
							0.319	0.405
Sargan (p-val)	0.918	0.877	0.856	0.72	0.635	0.635	0.808	0.696
m2 (p-val)	0.827	0.439	0.749	0.363	0.765	0.378	0.813	0.239
Time FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Obs	433	301	433	301	433	301	433	301
Countries	89	76	89	76	89	76	89	76

Note. The dependent variable is the 5-year log-difference of TFP level (da). All regressors are 5-year period averages. The sample spans between 1975 and 1999. All regressions include a constant. The estimation is performed with the two-step system-GMM procedure. Coefficients and standard errors are reported from the first step. *, ** and *** indicate that a coefficient is significant at 10, 5 and 1 per cent, respectively. The p-values for the Sargan overidentification test and the second order serial correlation (m2) test are reported from the second step.

Chapter 4

Explaining Co-movements Between Stock Markets: US and Germany*

1 Introduction

Measuring co-movements between stock markets is a widely debated issue. Since the seminal work of Grubel (1968), which expounded the benefits from international portfolio diversification, international stock markets co-movements have been analyzed in a series of studies. The early evidence of the seventies (see, for example, Granger and Morgenstern, 1970), despite divergent empirical methods used, led to the conclusions that correlations among returns to national stock markets are surprisingly low, and that national factors dominate their returns generating process. This results have been lately reversed by the empirical literature based on time series analysis, aimed at identyfiying separately trend and cycle components in equity markets. Kasa (1992) shows that equity markets in U.S., Japan, England, Germany and Canada over the period 1974-1990 share a single common stochastic trend, which mirrors the existence of a single common components in the structure of the dividend payments for all these markets. However, these results are not uncontroversial. Kanas (1998) analyzes daily data on US and major European stock markets to show that US market is not pairwise cointegrated with any of the European equity markets, hence there exist potential long-run benefits in risk reduction from diversifying

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in US stocks and European stocks. All these different results can be empirically reconciled by the literature which has shown that correlations between international equity markets vary strongly over time, and suggested two main distinct explanations for this phenomenon.[1] The first is based on the belief that the transmission mechanism is stable, while the features of shocks (global vs idiosyncratic) vary over time. In some periods global shocks do not occur and equity markets are driven by country-specific factors. As national business cycles are not well synchronized, all markets tend to move independently. In other periods all equity markets are globally affected by the same shocks and therefore their tendency to co-move increases. The alternative explanation relies upon the idea that periods of turbulence are characterized by the occurrence of shocks of unusual dimension, which may come along with structural breaks in their transmission mechanism. The empirical literature on the transmission of financial shocks (Rigobon, 1999) has recently formalized the distinction between the concepts of contagion and interdependence. The latter accounts for the existence of cross-market linkages, while contagion consists in modifications of such linkages during turbulent periods. Consider the case of the US and German stock markets: a strong co-movement of German and US equity prices in presence of unusual fluctuations in the US stock market is compatible both with interdependence and contagion. We have interdependence if the observed co-movement is in line with the historically measured simultaneous feedback between the two markets, while we have contagion when a change in the volatility of the US market (the disease) generates a structural break in the parameters measuring interdependence between US and German markets.

Identifying contagion from interdependence has important implications on the understanding of potential benefits from international portfolio diversification (see for instance, Rigobon and Forbes, 2002). In fact, in a world in which contagion is empirically relevant, optimal asset allocation should be regime dependent.The results on the benefits of diversification in a period of little turbulence are dramatically different from those in a period featuring large shocks.

Correlation between stock markets has been traditionally used when measuring co-movements and defining contagion. The earliest studies by King and Wadhwani (1990) and Bertero and Mayer (1990) presented and discussed the evidence of changes in unconditional covariances and correlations between stock returns on high-frequency data around the october 1987 crash. Since then, many authors pro-

[1] See Forbes and Rigobon (1998), Karoly and Stultz (1996), Lee and Kim (1993), Lin, Engle and Ito (1994), Longin and Solnik (1995, 2000).

posed different ways of testing the stability of (conditional) correlations, such as using ARCH and GARCH models (see Longin and Solnik, 1995 and Edwards and Susmel, 2000), cointegration (again, Longin and Solnik, 1995, Kasa, 1992, Serletis and King, 1997), or switching regimes (see Hassler, 1995 and Edwards and Susmel, 2000). This traditional approach has been recently criticized by Rigobon and Forbes (2002). It is easily shown that in a structural model featuring constant interdependence across countries, cross-market correlations are bound to increase in a period of turmoil, when stock market volatility increases. Hence, the evidence of changing patterns of correlations cannot be used to directly test for contagion. Rigobon and Forbes consider the 1997 East Asian crisis, the 1994 Mexican Peso crisis and the 1987 US stock market crash to show that unadjusted correlation coefficients support the contagion hypothesis, while tests based on coefficients adjusted for interdependence find virtually no-contagion. Alternative ways of correcting tests on correlations have been suggested, amongst the others, by Boyer et al. (1999) and Loretan and English (2000), that rely on normality of stock returns, and by Longin and Solnik (2001), who apply extreme value theory to conditional correlation coefficients and generalize their results for a wide class of returns distributions.[2]

An innovative methodology to test for contagion in presence of interdependence has been proposed by Rigobon (1999) through the implementation of an IV procedure. This strand of research crucially hinges on structural modelling of interdependence, with the adoption of a limited information approach.

This paper extends the limited information approach and test the hypothesis of "no contagion, only interdependence" through the full information estimation of a small co-integrated structural model, built following the LSE econometric methodology (see Hendry, 1995). Our measure of co-movements distinguishes between long-run and short-run dynamics for equity prices on different markets.

We concentrate on US and German stock markets, and consider a sample of monthly data spanning from January 1980 to September 2002. As a first step, we estimate a general reduced form VAR model on six variables (US and German share prices, earnings, and redemption yields on 10-year benchmark bonds). We remove non-normality and heteroscedasticity from the residuals by including a number of point dummies in the specification. Having obtained a valid specification for the VAR, we perform cointegration to identify long-run equilibria among the selected variables and attribute an exogeneity status to four of them (earnings and long-

[2] See Corsetti et al. (2002) for a survey and further contributions along this line of research.

term interest rates). Subsequently, we formulate a bivariate Vector Error Correction model for the two endogenous variables, i.e. equity prices in US and Germany. Finally, we proceed to specify a structural model of interdependence and test for no contagion. Our structural model is identified by assuming a lower triangular pattern of simultaneous feedbacks between US and German stock markets. On this model we test the further restrictions implied by the null of no contagion. Two orders of reasons led us to the choice of the German market as a representative for Europe at large. First, the German bond market has been a benchmark for the EMS over almost the entire sample, both in terms of volume and from a policy perspective. Since bond yields significantly enter our VAR and VECM specifications, it is more appropriate comparing the US with the major European market. Second, the German economy is a small open economy and economic fundamentals in Germany are dominated by an international trend, consistently with our lower triangular pattern of simultaneous feedback between US and German stock markets.

The paper is organized as follows. Section 2 explains our general-to-specific full-information approach to test for contagion, and compares it to alternative methodologies proposed by Forbes and Rigobon (2002), and by Rigobon (1999). Section 3 illustrates our empirical specification and contains a discussion of our analysis of long-run interdependence based on cointegration. Section 4 considers the short-run dynamics and illustrates how we attribute co-movements to interdependence and contagion. Section 5 concludes.

2 Estimating interdependence and contagion with small structural models

We consider the consensus definition of contagion as a change in the international propagation of shocks caused by some country specific factor. In the recent empirical literature on the international propagation of shocks such factor is usually interpreted as a crisis, identified by a local shock of different magnitude (usually paired with a change in the volatility of shocks). Measuring contagion requires some (structural) estimate of the mechanism of international propagation of shocks and the identification of a crisis.

To achieve this purpose we start from a reduced form VAR specification for the logarithms of US and German share prices, $LP_{Ger,t}, LP_{US,t}$ and the vectors of variables candidate to determine their equilibrium: $\mathbf{X}_{Ger,t}$, $\mathbf{X}_{US,t}$. For the sake of exposition, we consider a first order process, although our empirical model features higher order dynamics.

$$\begin{pmatrix} LP_{Ger,t} \\ LP_{US,t} \\ \mathbf{X}_{Ger,t} \\ \mathbf{X}_{US,t} \end{pmatrix} = \begin{pmatrix} \pi_{11} & \pi_{12} & \boldsymbol{\pi}'_{13} & \boldsymbol{\pi}'_{14} \\ \pi_{21} & \pi_{22} & \boldsymbol{\pi}'_{23} & \boldsymbol{\pi}'_{24} \\ \pi_{31} & \pi_{32} & \boldsymbol{\pi}_{33} & \boldsymbol{\pi}_{34} \\ \pi_{41} & \pi_{42} & \boldsymbol{\pi}_{43} & \boldsymbol{\pi}_{44} \end{pmatrix} \begin{pmatrix} LP_{Ger,t-1} \\ LP_{US,t-1} \\ \mathbf{X}_{Ger,t-1} \\ \mathbf{X}_{US,t-1} \end{pmatrix} + \begin{pmatrix} v_{1,t} \\ v_{2,t} \\ \mathbf{v}_{3,t} \\ \mathbf{v}_{4,t} \end{pmatrix}$$

$$\begin{pmatrix} v_{1,t} & \\ v_{2,t} & \\ \mathbf{v}_{3,t} & \mid I_{t-1} \\ \mathbf{v}_{4,t} & \end{pmatrix} \sim \left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \Sigma_t \right] \tag{4.1}$$

Note that residuals from our baseline VAR specification are heteroscedastic. This reflects the presence in the data of observations which correspond to periods of turmoil. By using tests of normality and heteroscedasticity of residuals as a guiding criterion, it is then possible to re-specify (4.1) as :

$$\begin{pmatrix} LP_{Ger,t} \\ LP_{US,t} \\ \mathbf{X}_{Ger,t} \\ \mathbf{X}_{US,t} \end{pmatrix} = \begin{pmatrix} \pi_{11} & \pi_{12} & \boldsymbol{\pi}'_{13} & \boldsymbol{\pi}'_{14} \\ \pi_{21} & \pi_{22} & \boldsymbol{\pi}'_{23} & \boldsymbol{\pi}'_{24} \\ \pi_{31} & \pi_{32} & \boldsymbol{\pi}_{33} & \boldsymbol{\pi}_{34} \\ \pi_{41} & \pi_{42} & \boldsymbol{\pi}_{43} & \boldsymbol{\pi}_{44} \end{pmatrix} \begin{pmatrix} LP_{Ger,t-1} \\ LP_{US,t-1} \\ \mathbf{X}_{Ger,t-1} \\ \mathbf{X}_{US,t-1} \end{pmatrix} \tag{4.2}$$

$$+ (I + \Psi D) \begin{pmatrix} u_{1,t} \\ u_{2,t} \\ \mathbf{u}_{3,t} \\ \mathbf{u}_{4,t} \end{pmatrix}$$

$$\begin{pmatrix} u_{1,t} & \\ u_{2,t} & \\ \mathbf{u}_{3,t} & \mid I_{t-1} \\ \mathbf{u}_{4,t} & \end{pmatrix} \sim N \left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \Sigma \right]$$

$$\Psi = \begin{pmatrix} \psi_{11} & \psi_{12} & \boldsymbol{\psi}'_{13} & \boldsymbol{\psi}'_{14} \\ \psi_{21} & \psi_{22} & \boldsymbol{\psi}'_{23} & \boldsymbol{\psi}'_{24} \\ \psi_{31} & \psi_{32} & \boldsymbol{\psi}'_{33} & \boldsymbol{\psi}'_{34} \\ \psi_{41} & \psi_{42} & \boldsymbol{\psi}'_{43} & \boldsymbol{\psi}'_{44} \end{pmatrix}$$

$$D = \begin{pmatrix} d_{1,t} & 0 & 0 & 0 \\ 0 & d_{2,t} & 0 & 0 \\ 0 & 0 & \mathbf{d}_{3,t} & 0 \\ 0 & 0 & 0 & \mathbf{d}_{4,t} \end{pmatrix}$$

where the vectors of dummies $\mathbf{d}_{i,t}$ are identified in order to filter non-normality out of the original residuals. The coefficients in the matrix Ψ allow the removal of outliers.

On the basis of this specification we proceed to cointegration analysis and reparameterise our system as follows:

$$\begin{pmatrix} \Delta LP_{Ger,t} \\ \Delta LP_{US,t} \\ \Delta \mathbf{X}_{Ger,t} \\ \Delta \mathbf{X}_{US,t} \end{pmatrix} = \mathbf{\Pi} \begin{pmatrix} LP_{Ger,t-1} \\ LP_{US,t-1} \\ \mathbf{X}_{Ger,t-1} \\ \mathbf{X}_{US,t-1} \end{pmatrix} + (I + \Psi D) \begin{pmatrix} u_{1,t} \\ u_{2,t} \\ \mathbf{u}_{3,t} \\ \mathbf{u}_{4,t} \end{pmatrix}$$

where the matrix $\mathbf{\Pi}$ describes the long-run properties of the system. In case of cointegration, there exist stationary combinations of the non-stationary variables. The rank of Π is reduced and equal to the number of cointegrating relationships, and we have $\mathbf{\Pi} = \boldsymbol{\alpha\beta}'$. The parameters in $\boldsymbol{\beta}$ describe the long-run equilibria of the system and by analyzing them we are able to address the issue of long-run interdependence. The parameters in $\boldsymbol{\alpha}$ describe the short-run response of the system to disequilibria and by analyzing them we are able to attribute the status of (weak) exogeneity to those variables that do not react to disequilibria. If weak exogeneity applies to the $\mathbf{X}$ variables and there is a unique cointegrating vector, we can simplify our general reduced form model in the following Vector Error Correction specification:

$$\begin{aligned} \begin{pmatrix} \Delta LP_{Ger,t} \\ \Delta LP_{US,t} \end{pmatrix} &= \begin{pmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \end{pmatrix} \boldsymbol{\beta}' \begin{pmatrix} LP_{US,t-1} \\ LP_{Ger,t-1} \\ \mathbf{X}_{Ger,t-1} \\ \mathbf{X}_{Ger,t-1} \end{pmatrix} \\ &+ \Pi_1 \begin{pmatrix} \Delta \mathbf{X}_{Ger,t} \\ \Delta \mathbf{X}_{US,t} \end{pmatrix} \\ &+ \left(I + \begin{pmatrix} \psi_{11} & \psi_{12} \\ 0 & \psi_{22} \end{pmatrix} \begin{pmatrix} d_{1,t} & 0 \\ 0 & d_{2,t} \end{pmatrix} \right) \begin{pmatrix} u_{1,t} \\ u_{2,t} \end{pmatrix} \end{aligned} \tag{4.3}$$

Note that the variables contained in the $\mathbf{X}$ vectors are now validly considered as exogenous. Moreover, the specification of the matrix Ψ is designed to match the empirical evidence that there are some German dummy variables that are not significant in the equation for US share prices while the converse is not true. The methodology can be extended to more general specifications for the vector of dummies (see for example Favero and Giavazzi, 2002).

The simultaneous presence of dummies in both equations is not informative on the relative importance of contagion and interdependence. This issue cannot be resolved by estimating a reduced form and requires the specification of a structural model. The following structural model, consistent with the reduced form (4.3), allows for both contagion and interdependence:

$$\begin{pmatrix} 1 & -\beta_{12} \\ 0 & 1 \end{pmatrix} \begin{pmatrix} \Delta LP_{Ger,t} \\ \Delta LP_{US,t} \end{pmatrix} = \begin{pmatrix} \gamma_{11} & \gamma_{12} \\ \gamma_{21} & \gamma_{22} \end{pmatrix} \boldsymbol{\beta}' \begin{pmatrix} LP_{US,t-1} \\ LP_{Ger,t-1} \\ \mathbf{X}_{Ger,t-1} \\ \mathbf{X}_{Ger,t-1} \end{pmatrix} + \tag{4.4}$$

$$+\,\Gamma_2 \begin{pmatrix} \Delta \mathbf{X}_{Ger,t} \\ \Delta \mathbf{X}_{US,t} \end{pmatrix} +$$

$$\left(I + \begin{pmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{pmatrix} \begin{pmatrix} d_{1,t} & 0 \\ 0 & d_{2,t} \end{pmatrix} \right) \begin{pmatrix} \epsilon_{1,t} \\ \epsilon_{2,t} \end{pmatrix}$$

$$\begin{pmatrix} \epsilon_{1,t} \\ \epsilon_{2,t} \end{pmatrix} \mid I_{t-1} \Bigg) \sim N \left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma^2_{\epsilon 1} & 0 \\ 0 & \sigma^2_{\epsilon 2} \end{pmatrix} \right]$$

In (4.4), we assumed triangularity in the simultaneous relationship between US and German stock prices, with the latter being influenced by the former but not vice-versa. This assumption, that characterizes our main identifying restrictions, is in line with the view that the US stock market has been playing a leading role amongst world markets, and is indeed supported by specific evidence in Eun and Shim (1989) and Cheung and Westermann (2001).[3] Note that in our empirical work we shall impose further restrictions on Γ_2, whose validity is testable as they are over-identifying restrictions. The presence of contagion is described by $a_{12} \neq 0$,

[3] Eun and Shim (1989) show, with a VAR approach, that innovations in the US stock market rapidly spread to a number of other national markets, including the German, while no single foreign market can explain the US market movements. Cheung and Westermann (2001) find that, on high frequency data, the lagged US equity returns are able to explain movements in the German indices, while the opposite is not true.

because this indicates that modelling interdependence by explicitly allowing $\beta_{12} \neq 0$ is not enough to describe the way shocks are transmitted across countries in periods of turmoil.

The null hypothesis of no contagion can then be tested as an over-identifying restriction for our specification. In particular, the hypothesis of interdependence only and no contagion is parametrized as $\mathrm{H}_0 : a_{12} = 0$, which implies the following overidentifying restriction:

$$\psi_{12} = \beta_{12} a_{22}$$

Under H_0, turmoil in country 2 propagates to country 1 only through interdependence, as described by β_{12}.

As extensively discussed in Favero and Giavazzi (2002), our full-information approach to test for contagion can be compared to the limited information approach, based on the IV method proposed by Rigobon (2000) to estimate β_{12} and control for interdependence in order to detect contagion. Rigobon's methodology hinges on splitting the sample into high and low volatility periods. Based on this distinction, an instrument is constructed whose validity is warranted under the null of no contagion, then tests of validity of instruments are used as a test of contagion. The beauty of this approach depends on the fact that it does not require variables other than endogenous to implement the IV estimator. In fact the instruments are constructed by taking transformation of the endogenous variables based on the presence of different regimes in volatility. Avoiding the estimation of a structural model of interdependence has the obvious benefit of imposing milder identifying restrictions than those necessary to implement our full-information procedure. The limited information approach has the advantage of identifying the system even when the traditional just-identifying restrictions are not valid. The main limit is that it is less powerful. The loss of efficiency could be non-negligible in cases where the number of observations for one of the two alternative regimes is low. Think of the limiting case in which the high-volatility sub-sample consists of very few observations: asymptotic results along the dimension of the full sample size are still applicable while obviously none applies along the dimension of the high volatility sub-sample. This is not a problem when daily or intra-daily high-frequency data are considered. However, it might become a problem when the potential importance of the role of fundamentals calls for the use of lower frequency data. In such a situation our methodology, based on a full information estimation on the whole sample with the inclusion of dummies for the high-volatility periods, is still applicable. Obviously, when the size of the high-volatility and low-volatility sub-samples are sufficiently long and the just-identifying restrictions in the structural model are validly imposed, the limited and full information approaches should both produce consistent estimators, and therefore the same results.

3 A statistical model for German and US share prices, earnings and long-term interest rates

Our statistical analysis of the relevance of contagion hinges on modelling both short-run and long-run interdependence between stock markets. We model long-run interdependence via cointegration and short-run interdependence via a small simultaneous structural model.

To investigate more closely the nature of the possible long-run equilibria, we consider the following VAR specification as our baseline statistical model:

$$\begin{pmatrix} LP_{US,t} \\ LE_{US,t} \\ R_{US,t} \\ LP_{Ger,t} \\ LE_{Ger,t} \\ R_{Ger,t} \end{pmatrix} = \mathbf{A}_0 + \sum_{i=1}^{4} \mathbf{A}_i \begin{pmatrix} LP_{US,t-i} \\ LE_{US,t-i} \\ R_{US,t-i} \\ LP_{Ger,t-i} \\ LE_{Ger,t-i} \\ R_{Ger,t-i} \end{pmatrix} + \begin{pmatrix} e_{1t} \\ e_{2t} \\ e_{3t} \\ e_{4t} \\ e_{5t} \\ e_{6t} \end{pmatrix}, \tag{4.5}$$

where LP_{US} and LP_{Ger} are the logs of the share price indexes, LE_{US} and LE_{Ger} the logs of I/B/E/S analysts forecasts of earnings, R_{US} and R_{Ger} the yields to maturity of ten-year benchmark bonds for US and Germany. Some discussion of our choice of variables and lag specification is in order.

Our choice of variables allows us to evaluate a number of different hypotheses recently adopted in the literature for the specification of long-run equilibria. Recent studies (Lander et al., 1997), following the time honoured contribution by Graham and Dodd (1962), have chosen to construct an equilibrium for stock markets by concentrating on long-term interest rates and the earning-price ratio. The long-term interest rate features a much stronger co-movement with price-earning ratios than the short-term interest rate. Such evidence can be rationalized by considering that the long-term interest rates contain an element of risk premium which is absent in the short-term interest rates. Studies concentrating on the relationship between the short-term interest rates and dividend or earning yields have found empirical evidence of a sizeable and strongly persistent risk premium (see Blanchard, 1983 and Wadhwani, 1998), which induces a rather weak long-run relationship between these variables.

Kasa (1992) has applied cointegration analysis to find a single common stochastic trend (and hence four cointegrating vectors) among the G5 stock market indexes. Serletis and King (1997) perform cointegration analysis in a framework similar to that of Kasa (1992) on ten EU stock markets. They measure the degree of convergence by applying time-varying parameter techniques to the vector of loadings measuring the short-run response of variables to disequilibria with respect to the cointegrating relationship(s). Our six-variables VAR allows to test the validity of the alternative long-run equilibria proposed by these authors on our data sets. Moreover, we can investigate the importance of long-run interdependence between stock markets by evaluating the relative importance of domestic and international factors in the determination of long-run equilibria.

Turning to the data, some graphical evidence on a sample of monthly data over the period 1980-2002 is provided in Figures 1-3, where we report yields to maturity on 10-year German and US Treasury bonds along with the (log) of earning/price ratio for the US and German stock markets.[4]

Insert Figure 1-3 here

The time-series behaviour of the reported variables suggests that long-term interest rates and US price-earning ratios might share a common stochastic trend while the existence of such common trend is more dubious for the German case, in which deviations from the trend tend to be more pronounced and more persistent. Very little evidence in favour of the hypothesis of common international stochastic trends in stock markets seems to emerge from our data.

Turning to the lag selection and the VAR specification, we have chosen the length of the distributed lags relying on the traditional likelihood based criteria. Note that, when we test for normality, heteroscedasticity and autocorrelation, strong evidence of non-normality emerges. Table 1 reports tests of the null hypothesis of residuals

[4] Our data-set comes from DATASTREAM. The stock price indexes are the Datastream all market indexes for US and Germany, the price earning ratios are from the same source and they are based on expected I/B/E/S analysts forecasts for end-of-year earnings. Finally, we considered yield-to-maturity for 10-year benchmark Treasury bonds. All data and an exact description of the Datastream stock market indexes are available from the website http://www.igier.uni-bocconi.it/personal/favero/homepage.htm

normality, both at the single equation and at the system level, proposed by Doornik and Hansen (1994).

Insert Table 1 here

The null of normality is rejected at the one per cent confidence level for all equations in the system. As a consequence, also normality of the vector of VAR residuals is strongly rejected. These diagnostic tests, which are in general important to detect misspecification and to ensure validity of inference, take additional importance in our context. In fact, non-normality is possibly determined by the presence of outliers, capturing the occurrence of those periods of turmoils that are crucial for detecting contagion. In order to ensure congruency of our statistical model and be able to exploit the information contained in the periods of turmoil, we proceed to include a number of point dummies in our specification. More precisely, we use an automatic criterion and construct a point dummy (taking a value 1 for the relevant observation and zero everywhere else) for each estimated residual lying outside the ± 2.5 standard deviation interval.[5]

As witnessed by the results reported in Table 2, the introduction of dummies largely solves the non-normality problems for all equations in our system, with the exception of equations for earnings, where non-normality is not attributable to specific large outliers, but to a consistent number of outliers of moderate dimension.

Insert Table 2 here

After controlling for outliers, we consider the following VAR as the baseline statistical model for our investigation:

$$\begin{pmatrix} LP_{US,t} \\ LE_{US,t} \\ R_{US,t} \\ LP_{Ger,t} \\ LE_{Ger,t} \\ R_{Ger,t} \end{pmatrix} = \mathbf{A}_0 + \sum_{i=1}^{4} \mathbf{A}_i \begin{pmatrix} LP_{US,t-i} \\ LE_{US,t-i} \\ R_{US,t-i} \\ LP_{Ger,t-i} \\ LE_{Ger,t-i} \\ R_{Ger,t-i} \end{pmatrix} + \mathbf{B} * \mathbf{DUM} + \begin{pmatrix} e_{1t} \\ e_{2t} \\ e_{3t} \\ e_{4t} \\ e_{5t} \\ e_{6t} \end{pmatrix}, \tag{4.6}$$

[5]The threshold has been chosen on the basis of the normality of residuals after the dummies have been included in the specification. Our results are robust to modification of such threshold in the range of 2-3 standard deviations

where **DUM** is a vector of thirty-three dummies, taking value of one when the outlier occurs and zero anywhere else.

Endowed with model (4.6), we address the first issue of our interest: long-run equilibrium and interdependence between US and German stock markets.

Re-parameterize (4.6) as follows:

$$\begin{pmatrix} \Delta LP_{US,t} \\ \Delta LE_{US,t} \\ \Delta R_{US,t} \\ \Delta LP_{Ger,t} \\ \Delta LE_{Ger,t} \\ \Delta R_{Ger,t} \end{pmatrix} = \mathbf{A}_0 + \sum_{i=1}^{3} \mathbf{\Pi}_i \begin{pmatrix} \Delta LP_{US,t-i} \\ \Delta LE_{US,t-i} \\ \Delta R_{US,t-i} \\ \Delta LP_{Ger,t-i} \\ \Delta LE_{Ger,t-i} \\ \Delta R_{Ger,t-i} \end{pmatrix} + \Pi \begin{pmatrix} LP_{US,t-1} \\ LE_{US,t-1} \\ R_{US,t-1} \\ LP_{Ger,t-1} \\ LE_{Ger,t-1} \\ R_{Ger,t-1} \end{pmatrix} +$$

$$+ \mathbf{B} * \mathbf{DUM} + \begin{pmatrix} e_{1t} \\ e_{2t} \\ e_{3t} \\ e_{4t} \\ e_{5t} \\ e_{6t} \end{pmatrix}, \tag{4.7}$$

$$\mathbf{\Pi}_i = -\left(I - \sum_{j=1}^{i} \mathbf{A}_j \right),$$

$$\mathbf{\Pi} = -\left(I - \sum_{i=1}^{3} \mathbf{A}_i \right),$$

where the matrix $\mathbf{\Pi}$ describes the long-run properties of our system. In particular, the rank of $\mathbf{\Pi}$ determines the number of cointegrating vectors. Whenever the rank of $\mathbf{\Pi}$ is reduced, the following decomposition applies $\mathbf{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$, where the matrix $\boldsymbol{\beta}$ contains the parameters in the cointegrating vector(s) and the matrix $\boldsymbol{\alpha}$ contains the loadings describing the adjustment of each variable to disequilibria with respect to the long-run equilibrium of the system. We analyze the rank of $\mathbf{\Pi}$ and its decomposition by using the statistical framework proposed by Johansen (1995).

Insert Table 3 here

Table 3 reports the sequence of estimated eigenvalues of the long-run matrix along with the test for the rank of Π based on the trace-statistic and the maximum eigenvalue statistics, which points toward the existence of a unique cointegrating vector.[6]

Having fixed the rank of $\mathbf{\Pi}$ to one, we test alternative hypotheses on the specification of the long-run relationship. We consider four alternative hypotheses. H_1 postulates a long-run relation between the log of US price-earning and yields to maturity of US and German long-term bonds. Under H_2 a long-run relation exists between the log of German price-earning and yields to maturity of US and German long-term bonds. H_3 claims that a long-run relation links the log of German price-earning ratio to the log of US price-earning ratio, and finally H_4 postulates a long-run relation between the log of German stock price and the log of US stock price. The first two hypotheses reflect a generalized version of the long-run solution based on Graham and Dodd and adopted by Lander et al. (1997): H_1 applies it to the US, while H_2 applies it to Germany. H_3 and H_4 allow explicitly for interdependence among US and German stock markets, following the specification of the cointegrating relations chosen by Kasa (1992) and Serletis and King (1997).

Only hypothesis H_1, implying a long-run relationship between the (log of) US price-earning ratio and the US and German long-term interest rates is not statistically rejected. Moreover, the loadings associated to the cointegrating vector show that only the US and German stock prices significantly react to disequilibrium. Hence earnings and long-term interest rates can be considered as weakly exogenous for the estimation of the parameters of interest when estimating models for share prices. Our cointegrating relationship is directly comparable with that obtained by Lander et al. (1997). In fact, we obtain very similar results except that the long-term interest rates relevant to our cointegrating vector are some weighted average of the US and German long-term rates. Figure 1 may help the interpretation of such results: over the second part of our sample there is virtually no difference between the two long-term rates, while in the first part of the sample the US long-term rates fluctuate remarkably more than the German ones. Price/earning ratios differ from the nominal long term interest rates in that they are real variables and hence they are less affected by inflation.[7] Cointegration between price/earnings and long-term

[6] We report both trace and maximum eigevalue statistics although there is evidence that the trace statistic is preferable as the sequence of trace tests lead to a consistent test procedure, while no such result is available for the max eigenvalue statisatics see (see Doornik and Hendry, 2001). The presence of dummies makes the traditional critical values not appropriate, although the difference of magnitude in the sequence of eigenvalues suggests that the evidence in favour of the existence of a unique cointegrating vector is robust.

[7] By inflation here we mean average ten-year inflation. In fact our price-earning ratios, being defined with reference to expected earnings, are indeed affected by short-term, one-period ahead,

nominal rates implies stationarity of inflation. Current analysis of U.S. monetary policy generally acknowledges that 1979 marks the beginning of a new policy regime characterized by a strong anti-inflationary stance which allowed a mean reverting relation between effective inflation and the target chosen by the monetary policy authorities.[8] Despite the change in the monetary policy regime, some episodes of "inflation scares" hit the US bond market at the beginning of the new monetary regime. As these episodes remained local, some weighted average of the US and German rates is not so dramatically affected by the temporary jumps in expected inflation and keeps a better balance with the price/earning ratio.

We conclude this section by reporting in Figure 4 the deviation of US share prices from their equilibrium value.

Insert Figure 4 here

The Figure shows twenty episodes of mean reversion over twenty years. It also suggests that the US market was heavily overvalued at the beginning in 1982 and at the end of year 2000, while at the end of September 2002 share prices fluctuated at thirty per cent discount with respect to their equilibrium value.

4 Measuring short-run interdependence and contagion

To describe short-run interdependence and assess contagion we need a structural model. We build it starting from simplifying the baseline statistical model into a bivariate Vector Error Correction model for US and German share prices, where, on the basis of the statistical evidence on the loadings reported in Table 3, long-term interest rates and earnings are taken as weakly exogenous :

inflation.

[8]Empirical investigations of the Fed's reaction function confirm this discontinuity. See the widely cited work of Clarida, Gali and Gertler (2000). Cogley and Sargent (2002) also relate the conquest of U.S. inflation to a different behaviour of the monetary policy authority under the Volcker and Greenspan tenures.

$$\begin{pmatrix} \Delta LP_{US,t} \\ \Delta LP_{Ger,t} \end{pmatrix} = \mathbf{B}_0 + \mathbf{B}_1 \begin{pmatrix} \Delta LP_{US,t-1} \\ \Delta LP_{Ger,t-1} \end{pmatrix} + \sum_{i=0}^{1} \mathbf{C}_i \begin{pmatrix} \Delta R_{US,t-i} \\ \Delta R_{Ger,t-i} \\ \Delta LE_{US,t-i} \\ \Delta LE_{Ger,t-i} \end{pmatrix} + \tag{4.8}$$

$$+ \mathbf{D}\left(LP_{US,t-1} - LP^{*}_{US,t-1}\right) + \mathbf{F}\left(dum\right) + \begin{pmatrix} u_{1t} \\ u_{2t} \end{pmatrix},$$

$$LP^{*}_{US,t-1} = LE_{US,t-1} - 0.10R_{US,t-1} - 0.111R_{Ger,t-1} + 4.4,$$

The vector of dummies dum is a sub-vector of the one containing thirty-three dummies used for the general system in the cointegration analysis. The dynamics of the system considering earnings and long-term rates as exogenous is much shorter, since the first order dynamics is now chosen through optimal lag selection criteria. There are twelve outliers, among which nine are common to both equations and three are specific to the equation for the German share price. The common dummies correspond to episodes of US stock market turmoil. In 1987:10, 1998:08, 2001:02, 2001:03, 2001:09 2002:7 and 2002:09, we observed downward movements respectively of twenty-four, twelve, eleven, nine, seven and a half and nine per cent, while in 1987:01 and 1998:10 equity prices jumped up by thirteen and twelve per cent. Country specific movements in German equity prices are accounted for by the dummies respectively of 1990:09 and 1997:08, and of 1999:12, when the market fell by nineteen and eleven and rose by thirteen per cent. The diagnostic tests reported in Table 4 show that the null of absence of residuals correlation, homoscedasticity and normality cannot be rejected for (4.8).[9]

Insert Table 4 here

On the basis of this reduced form, we proceed to estimate two structural models. As discussed in section 2, we consider a more general one allowing for both short-run interdependence and contagion, and a more restrictive model consistent with the hypothesis of "only interdependence, no contagion".[10] Both structural models

[9] All the tests have performed at system level using PC-FIML. For a detailed description see Doornik-Hendry (1997)

[10] Note that our discussion in section 2 introduces multiplicative dummies on the residuals, this is equivalent to the introduction of shift dummies in the structural model. In fact, with point dummies multiplicative effect are observationally equivalent to shift effects.

impose some testable over-identifying restrictions on our reduced form and we can therefore use the outcome of the tests to discriminate between the cases of interest. The estimated structural models are reported in Table 5. Both models show that the fluctuations in local fundamentals, such as earnings and the long-term interest rates, determine fluctuations in share prices. The US market also react very significantly to deviation of US share prices from their long-run equilibrium. Such variables also affect the fluctuations in German prices although the effect is quantitatively smaller and just marginally statistically significant. Model 1 in Table 5 is consistent with the hypothesis of the existence of contagion between the US and German stock markets. In fact, in the case of interdependence only, when a simultaneous feedback is allowed from US to European stock market, the dummies capturing turmoil periods in the US market should not enter significantly the equation for German stock prices. We observe that not only do such dummies enter significantly, but their inclusion also renders the simultaneous feedback between German and US stock markets not significantly different from zero. Importantly, the model is supported by the data in that the tests for the validity of the ten over-identifying restrictions imposed by Model 1 on the general reduced form (4.8) does not lead to the rejection of the null hypothesis of interest.

The results from the estimation of the structural model implicit in the hypothesis of "no contagion, only interdependence" are reported in the same Table under the label of Model 2. The validity of over-identifying restrictions is now rejected. As we have nine dummies for the US stock market, our test for the null of no-contagion is distributed as a χ^2_{29}, with nine more degrees of freedom than the statistic used to test the validity of Model 1. Interestingly, as a consequence of the omission of dummies, the significance of the simultaneous feedback increases drastically and might mislead the inference whenever Model 2 is estimated without reference to the general model (4.8).

Insert Tables 5

To allow comparison of our results with the IV based approach we have created an instrument w_t,which is equal to $-\Delta LP_{US,t}/261$ for all observations in our sample except for 1987:01 ,1987:10, 1998:08, 1998:10, 2001:02, 2001:03, 2001:09 and 2002:09, where it takes value $\Delta LP_{US,t}/9$. We report in Table 6 the results of the regression that show the validity of w_t as an instrument for $\Delta LP_{US,t}$. Here, we also present the augmented regression for the German share prices, that allows to implement the

Hausman- type test for the validity of instrument, suggested by Rigobon as a test for contagion.

Insert Table 6 here

As the coefficient on $\hat{u_t}$ is significantly different from zero, the null of no-contagion is rejected and our results are confirmed by the implementation of the IV procedure.

5 Conclusions

In this paper we have proposed a methodology to disentangle interdependence from contagion in co-movements between stock markets and applied it to the case of the German and US stock markets. We assessed the relative importance of contagion and interdependence within the framework of an explicit structural model, using cointegration analysis to separate long-run equilibria from short term dynamics. We constructed our long-run equilibria by tsting different possible specification and favouring the hypothesis of cointegration between the (log of) US earning-price ratio and long-term interest rates. Within such framework, we found that the hypothesis of no long-run interdependence between the two markets cannot be rejected. We then used our Vector Error Correction Model as a baseline reduced form and constructed a structural model to assess the relative importance of interdependence and contagion in determining the short-run dynamics of the two markets. Our structural model shows that the effect of fluctuations of US stock market on the German stock market is captured by a non-linear specification. Normal fluctuations in the US stock market have virtually no effect on the German market, while such effect becomes sizeable and significant for abnormal fluctuations. Such non-linearity is clearly consistent with the relevance of contagion, in that it amounts to a modification of short run interdependence in periods of turmoil. Our results are proven to be consistent with those obtained by applying the Instrumental Variable methodology proposed by Rigobon (1999). We believe that our findings have important implications for international portfolio diversification. In fact, our empirical evidence of no long-term interdependence between US and German stock markets speaks in favour of benefits from diversification of an asset allocation with a long-term horizon. On the other side, our empirical evidence on the importance of contagion in the short-term interdependence between the two markets illustrates the risk of any short-term asset allocation which does not explicitly recognizes the importance of non-linearity and regime-switching in the relation between international stock returns.

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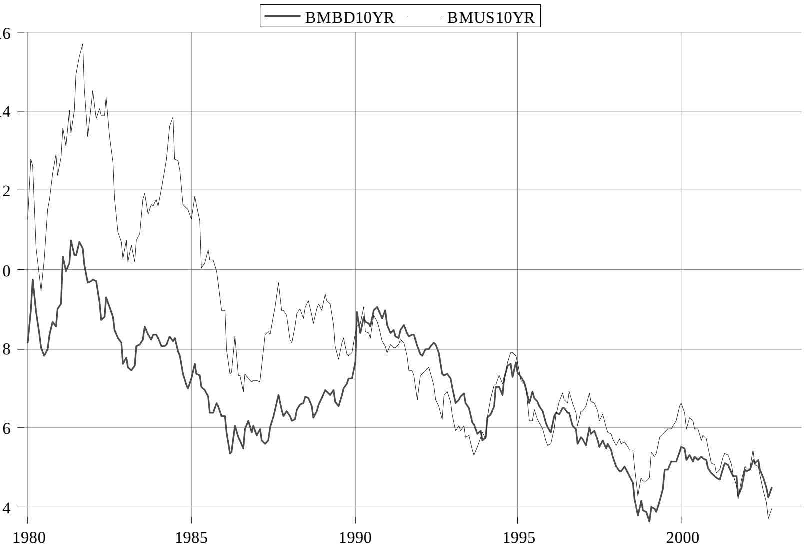

Figure 1: US (BMUS10YR) and German (BMBD10YR) long-term interest rates. Source: Datastream.

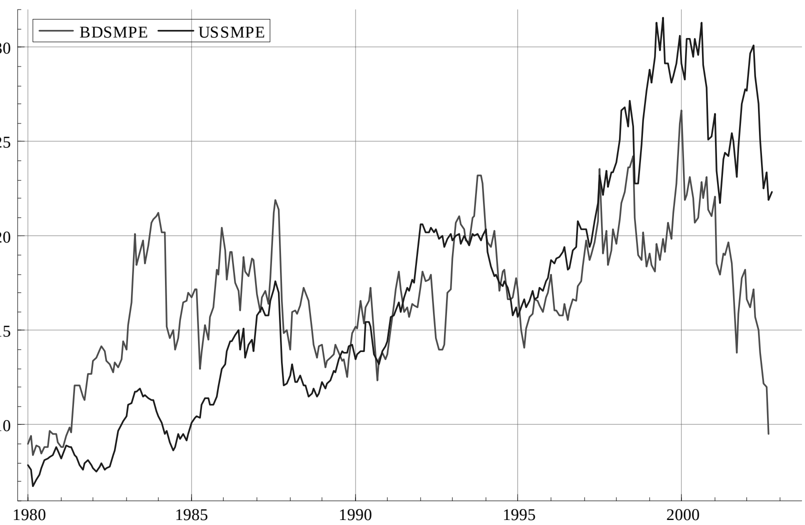

Figure 2: US (USSMPE) and German (BDSMPE) price/earning ratios. Source: Datastream.

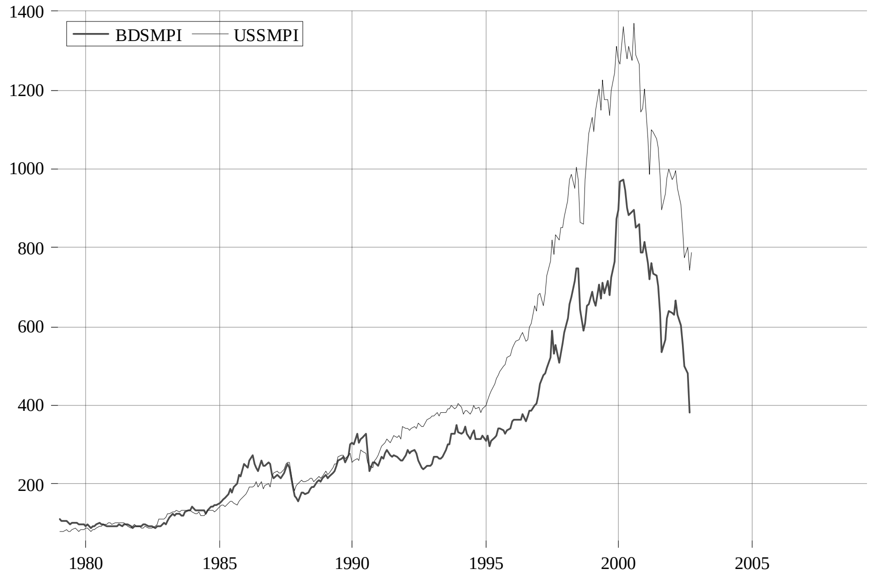

Figure 3: US(USSMPI) and German(BDSMPI) Datastream all market share price indexes

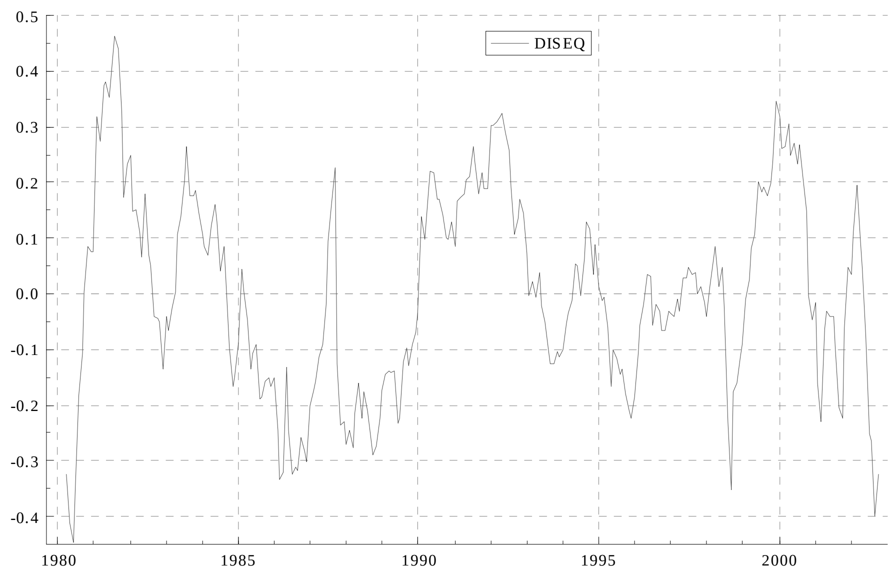

Figure 4: Deviation from long-run equilibrium of US share prices (0.x indicate a 10*x per cent deviation)

TABLE 1: Testing normality of residuals in the VAR system

Single equation	without dummies	with dummies
$LP_{US,t}$	35.951**	7.27*
$LE_{US,t}$	67.194**	21.22**
$R_{US,t}$	1.282	4.71
$LP_{Ger,t}$	34.276**	2.08
$LE_{Ger,t}$	83.044**	73.24**
$R_{Ger,t}$	34.502**	2.33

The estimated model is $\mathbf{Y}_t = \mathbf{A}_0 + \sum_{i=1}^{4} \mathbf{A}_i \mathbf{Y}_{t-i} + \mathbf{u}_t$, with

$Y_t = \left[LP_{US,t}, LE_{US,t}, R_{US,t}, LP_{Ger,t}, LE_{Ger,t}, R_{Ger,t}\right]'$,

dummies are introduced to eliminate outliers, defined as observed residuals with an absolute value larger than 2.5 time their standard deviation.

The test statistics reported are based on Hansen-Doornik (1994) and distributed as a χ^2_2 * and ** indicate rejection respectively at 5 and 1 per cent significance level.

TABLE 2: Testing the number of cointegrating vectors

Variable	$H_0 : rank\ \Pi = p$	Trace	Max Eig.
0.184	$p = 0$	110.3**	54.6**
0.066	$p \leq 1$	55.7	18.54
0.054	$p \leq 2$	37.13	14.98
0.045	$p \leq 3$	22.15	12.57
0.023	$p \leq 4$	9.58	6.43
0.01	$p \leq 5$	3.15	3.15

Eigenvalue column reports the estimated eigenvalues of Π.

Trace and Max Eig. columns reports the values of the trace and. maximum eigenvalue statistics for the null $rank\ \Pi = p$, i.e. there are at most p cointegrating vectors

TABLE 3: Testing hypothesis on the long-run equilibrium

	H_1		H_2	H_3	H_4
Variable	β	α	β	β	β
$LP_{US,t}$	1	-0.05 (0.015)	0	1	1
$LE_{US,t}$	-1	0.005 (0.006)	0	-1	-1
$R_{US,t}$	0.10 (0.016)	-0.13 (0.15)	-0.013 (0.032)	0	0
$LP_{Ger,t}$	0	-0.056 (0.018)	1	-6.33 (2.68)	0
$LE_{Ger,t}$	0	-0.039 (0.15)	-1	6.33 (2.68)	0
$R_{Ger,t}$	0.10 (0.03)	-0.17 (0.09)	-0.10 (0.058)	0	0
Test restrictions	χ^2_3= 10.78		χ^2_3= 30.92**	χ^2_4= 34.86**	χ^2_5= 36.57**

Tests on 4 alternative restrictions on the unique cointegrating vectors.

H_1 long-run relation between log US price-earning and R_{US} and R_{GER}.

H_2 long-run relation between log German price-earning and R_{US} and R_{GER}.

H_3: long-run relation between the logs of German and US price-earning ratios.

H_4: long-run relation between the logs of US and German stock prices.

For each hypothesis we report estimated parameters in the cointegrating vector, with the associated standard error and, whenever the validity of the over-identifying restrictions is not rejected, the estimated loadings of the cointegrating vectors in equations associated to each variable in the VAR.

TABLE 4: Testing congruency of the bivariate VECM

	Normality	Autocorrelation 1-7	Heteroscedasticity
$\Delta LP_{US,t}$	$\chi^2_2 = 11.1(0.01)$	$F_{7,237} = 0.8(0.63)$	$F_{46,204} = 1(0.43)$
$\Delta LP_{Ger,t}$	$\chi^2_2 = 3.6(0.17)$	$F_{7,237} = 0.6(0.76)$	$F_{46,204} = 1.3(0.11)$
System	$\chi^2_4 = 4.9(0.30)$	$F_{28,472} = 0.7(0.87)$	$F_{138,606} = 1.1(0.22)$

The table reports tests for Normality, Heteroscedasticity and Autocorrelation of the residuals for the following model:

$$\begin{pmatrix} \Delta LP_{US,t} \\ \Delta LP_{Ger,t} \end{pmatrix} = \mathbf{B}_0 + \mathbf{B}_1 \begin{pmatrix} \Delta LP_{US,t-1} \\ \Delta LP_{Ger,t-1} \end{pmatrix} + \sum_{i=0}^{1} \mathbf{C}_i \begin{pmatrix} \Delta R_{US,t-i} \\ \Delta R_{Ger,t-i} \\ \Delta LE_{US,t-i} \\ \Delta LE_{Ger,t-i} \end{pmatrix} +$$

$$+ \mathbf{D} \left(LP_{US,t-1} - LP^*_{US,t-1} \right) + \mathbf{F} \left(\mathbf{dum} \right) + \begin{pmatrix} u_{1t} \\ u_{2t} \end{pmatrix},$$

$$LP^*_{US,t-1} = LE_{US,t-1} - 0.10 R_{US,t-1} - 0.111 R_{Ger,t-1} + 4.4,$$

where the vector dum contains dummies for periods 87:1, 87:10, 90:9, 97:8, 98:8, 98:10, 99:12, –01:2, 01:3, 01:9, 02:9. Rows 2 and 3 report the relevant statistics for each equation with p-values in parentheses, while row four does the same for the system. Normality test, based on Hansen and Doornik (1994), is rejected for the first equation but neither for the second nor for the system. LM test for autocorrelation up to the seventh order and White test for heteroscedasticity of residuals are rejected both for single equations and for the entire system

TABLE 5: Structural models for US and European stock prices

	Model 1		Model 2	
	$\Delta LP_{US,t}$	$\Delta LP_{Ger,t}$	$\Delta LP_{US,t}$	$\Delta LP_{Ger,t}$
Constant	**0.009** (0.002)	**0.008** (0.004)	**0.006** (0.002)	−0.0007 (0.003)
$\Delta LP_{US,t}$		−0.14 (0.271)		**0.90** (0.109)
ECM_{t-1}	**−0.053** (0.012)	**−0.036** (0.021)	**−0.055** (0.012)	0.026 (0.017)
$\Delta R_{US,t}$	**−0.032** (0.005)		**−0.024** (0.005)	
$\Delta R_{Ger,t}$		**−0.048** (0.011)		−0.019 (0.010)
$\Delta LE_{US,t}$	**0.19** (0.094)		0.18 (0.111)	
$DUM8701$	**0.108** (0.035)	**−0.125** (0.056)	0.061 (0.034)	
$DUM8710$	**−0.259** (0.035)	**−0.29** (0.083)	**−0.264** (0.034)	
$DUM9009$		**−0.168** (0.038)		**−0.176** (0.044)
$DUM9708$		**−0.083** (0.039)		−0.082 (0.044)
$DUM9808$	**−0.141** (0.035)	**−0.19** (0.056)	**−0.150** (0.034)	
$DUM9810$	**0.104** (0.035)	0.047 (0.053)	**0.088** (0.034)	
$DUM9912$		**0.104** (0.039)		**0.112** (0.044)
$DUM0102$	**−0.126** (0.035)	−0.095 (0.054)	**−0.120** (0.034)	
$DUM0103$	**−0.101** (0.035)	−0.09 (0.056)	**−0.098** (0.034)	
$DUM0109$	**−0.112** (0.035)	**−0.19** (0.056)	**−0.130** (0.034)	
$DUM0207$	**−0.11** (0.033)	**−0.14** (0.056)	**−0.141** (0.034)	
$DUM0209$	**−0.107** (0.035)	**−0.28** (0.056)	**−0.141** (0.034)	
LR test	$\chi^2_{19} = 20.73(0.37)$		$\chi^2_{28} = 101.94(0.000)$	

Model 1 reflects the hypothesis of "interdependence and contagion" among US and German stock markets. Model 2 reflects the hypothesis of no contagion. The LR test is a statistic for the valifity of the over-identifying restrictions imposed by each model on the Vector Error Correction Reduced form reported in Table 4

TABLE 6: Testing contagion by the limited information approach

	$\Delta LP_{US,t}$	$\Delta LP_{Ger,t}$
w_t	**0.010** (0.002)	
Constant	**7.884** (0.877)	−0.007 (0.005)
ECM_{t-1}		−0.030 (0.025)
$\Delta LE_{Ger,t}$		0.281 (0.463)
ΔR_{Ger}		**−0.056** (0.019)
$\Delta LP_{US,t}$		**1.137** (0.181)
$\widehat{u}_t$		**−1.488** (0.469)
$DUM9009$		**−0.195** (0.057)
$DUM9708$		−0.154 (0.081)
$DUM9912$		**0.164** (0.061)

Column 2 reports OLS coefficients from the regression of $\Delta LP_{US,t}$ on w_t, the instrument suggested by Rigobon, which equals $\frac{\Delta LP_{US,t}}{9}$ for the observations corresponding to US-specific dummies and $-\frac{\Delta LP_{US,t}}{264}$ elsewhere. Colum 3 reports coeffcients estimated by IV. Instruments are: $\Delta LE_{US,t}, \Delta R_{US,t}, \Delta LE_{Ger,t}, \Delta R_{Ger,t}, ECM_{t-1}$,DUM8701, DUM8710, DUM9009, DUM9708, DUM9808, DUM9810, DUM9912, DUM0102 DUM0103, DUM0109, DUM0207, DUM0209. u_t are estimated residuals from the regression in column2. Standard errors are reported in parentheses and significant coefficients in bold.

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