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Legal Determinants of the Return on Equity

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Davide Lombardo[*] and Marco Pagano[**]

Abstract

Recent work documents that better legal institutions are associated with broader equity markets. We investigate whether differences in legal institutions also help explain the international cross-section of expected stock returns. Three main regularities emerge. First, total stock market returns are positively correlated with overall measures of the quality of institutions, such as judicial efficiency and rule of law, controlling for risk. Second, dividend yields and earning-price ratios also correlate positively with judicial efficiency and rule of law, controlling for risk and expected earnings growth. Thirdly, equity returns have no relationship with the degree of protection of minority shareholders granted by the letter of the law. These findings suggest that equity markets are segmented, and that differences in monitoring, auditing and other enforcement costs are not the sole reason for segmentation. We interpret the positive cross-country correlation between the overall quality of the legal system and the expected return on equity as resulting from the curtailment of insiders' private benefits and the increase of firms' profitability associated with better institutions.

Keywords: law, enforcement, shareholder protection, corporate governance, return on equity

JEL Classification G12, K22, K42

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1. Introduction

Understanding the determinants of the international cross-section of stock returns has proved a daunting task for research to date. The explanatory power of the International Capital Asset Pricing Model (ICAPM) is typically found to be quite low (see, among others, Solnik, 1977, Ferson and Harvey, 1994). Several authors (see Bekaert and Harvey, 1995, for example) attribute this to the empirical failure of the assumption that equity markets are perfectly integrated. In this paper we suggest that a better understanding of this subject can be achieved by combining the traditional asset pricing approach with the law and finance approach – two strands of literature so far totally disjoint.

Recent cross-country empirical work by La Porta, Lopez de Silanes, Shleifer and Vishny (1997, 1998) [henceforth LLSV] shows that external equity financing correlates with differences in legal rules and in their enforcement. More specifically, the share of stock market capitalization held by minorities and the number of initial public offerings correlate positively with the degree of respect for the law, the quality of judicial enforcement and the legal protection of minority shareholders' property rights. LLSV interpret this finding as evidence that opportunistic behavior by managers and controlling shareholders is less likely when the legal protection of investors is strong and its judicial enforcement is swift and cost-effective.[1]

An obvious question then arises. Do differences in the protection of minority shareholders' rights and, more generally, in the protection of property rights result also in cross-country differentials in risk-adjusted expected returns? In this paper, we take this question to the data.

In theory, the legal environment can affect the severity of agency problems between company insiders and outside shareholders in two ways. First, it may directly affect the private benefits that managers are able to extract from their companies. Second, it determines the auditing and judicial costs that shareholders expect to incur in order to secure payment of dividends by managers. International differences along these two dimensions of the legal environment have different effects on equilibrium stock returns, depending on the degree of international integration of stock markets.

In a perfectly integrated stock market, risk-adjusted equity returns should be independent of the legal system if this affects only the amount of corporate resources that managers can divert to themselves: equity prices will discount the private benefits to be extracted by managers, so as to equalize the expected risk-adjusted rate of return across countries. If instead, as a result of poor legal protection, investors must bear monitoring and auditing costs to enforce their claims to corporate dividends, equilibrium expected returns must compensate them for such costs, as well as for non-diversifiable market risk. In the latter case, more

[1] Modigliani and Perotti (1997) also highlight the relationship between the enforceability of contracts and the availability of external finance.

investor-friendly legal institutions should be associated with lower risk-adjusted returns on equity.[2]

When international equity markets are segmented for additional reasons (see for instance Gordon and Bovenberg, 1996), the relationship between risk-adjusted returns and the quality of the institutional framework is more complex. First, with segmented equity markets, also the amount of private benefits matters for equilibrium stock returns. Since the risk-bearing capacity of domestic investors is limited, higher amounts of equity funding come forth only if companies are willing to pay a higher expected rate of return. Stricter legal limits to managerial opportunism allow companies to credibly pledge higher returns and thereby obtain a larger amount of equity funding. In this case, therefore, countries with lower levels of private benefits of control can feature higher risk-adjusted rates of return, in equilibrium.

Under segmented markets, institutions can also affect equilibrium returns through a completely different channel, independently of the presence of agency problems between insiders and outside investors. Namely, fair legal rules and a trustworthy judicial system may widen the menu of enforceable contracts, thus increasing the profitability of the corporate sector and hence its demand for (equity) financing. Ceteris paribus, also this effect would result in a higher equilibrium rate of return on equity.

In our study, we face some well-known problems in measuring risk-adjusted expected returns on equity. The first is that expectations are not directly observable. Second, several measures of the return on equity can be used. For a public company, the return on its shares, the earnings-price ratio, and the dividend yield are all candidate measures, each with its own strengths and weaknesses. For a private company seeking its first listing, instead, a more relevant measure may be the excess return of its shares in initial public offerings (IPOs). Thirdly, measures such as the return on secondary markets must be corrected for undiversifiable risk (and for diversifiable risk if equity markets are internationally segmented), while others, such as the earnings-price ratio, must be purged of the component reflecting cross-country differentials in expected earnings growth.

We document a rich set of empirical regularities. Of these, the most robust finding is that the risk-adjusted return on equity is positively associated with the respect for the law and the efficiency of the judicial system. This is evidence that equity markets are segmented, since under capital market integration this correlation should be zero, or even negative if cross-country legal differences translate into different auditing and monitoring costs for investors. This finding agrees with much of the empirical literature in this area (see for example Roll (1992), Heston and Rouwenhorst (1994) and Bekaert and Harvey (1995), among others). The novelty of our evidence is that, even controlling for cross-country differences in legal and judicial institutions, there are other sources of international segmentation affecting the cross-section of expected returns. We also find that popular measures of the formal protection of minority shareholders' rights are not correlated with risk-adjusted equity returns.

[2] A similar effect arises in Merton's (1987) CAPM model when an innovation reduces the costs of obtaining reliable information about a company and thereby expands its shareholder base. The expected return on the stock declines as the fraction of investors informed about the stock rises.

Our paper also contributes to the small but growing empirical literature on law and finance, and especially to the interpretation of the seminal work by LLSV.[3] First, our results show the importance of controlling for risk in assessing the correlation between institutional variables and asset returns. LLSV (1999), using international company-level data, find that the cash-flow/price ratio is negatively related to shareholder rights protection, and interpret this as resulting from less severe agency problems. But their correlation could be driven by an omitted risk variable. If, for example, countries with better protection of shareholder rights are also safer investment havens for an international asset manager, then companies in those countries would fetch higher valuation irrespective of agency problems.[4] Second, our evidence of international equity market segmentation suggests that also the results of LLSV (1997) should be taken with caution: the relative size of equity markets may be affected by omitted country-specific factors, over and beyond the quality of each country's legal institutions, potentially leading to biased estimates. Using a panel of industry-level data, Lombardo (2000) builds on this point, and controls for country effects (and industry effects) in his empirical assessment of the effect of institutions on the risk-adjusted expected return on equity.

The structure of the paper is as follows. Section 2 presents a simple analytical framework to interpret the evidence. In section 3 we describe the data. In section 4 we report our results, which are based on different measures of the return on equity: secondary market returns, accounting measures, and excess returns on primary equity issues. Section 5 concludes.

2. A Simple Model

In this section we briefly illustrate the channels through which legal variables may affect the equilibrium rate of return on equity. The analysis, for brevity presented only graphically, is based on a simple model presented in Lombardo and Pagano (2000), and is meant as a tool to organize ideas and interpret the empirical findings presented in subsequent sections.

[3] Our paper is indirectly related to a set of recent papers on the effect of financial development on growth, which employ legal variables mainly to instrument financial development measures in their growth regressions. Rajan and Zingales (1998) point out that firms and industries that are more dependent on external finance tend to grow faster in countries where financial markets are better developed, and test this prediction on a large panel of industry-level cross-country data. Carlin and Mayer (1998) build on the Rajan-Zingales approach to probe further into the relationships between industrial activity, financial systems and legal arrangements, and conclude that market-based finance and legal protection of investors are correlated with the growth of equity-financed and skill-intensive industries. Demirgüc-Kunt and Maksimovic (1998) test the same hypothesis on firm-level data from thirty countries. They estimate the maximum growth rate that each firm of their sample could attain without access to long-term financing, and compare these potential growth rates to those attained by firms in countries with different legal and financial systems. They show that in countries with better legal systems, more active stock markets and larger banking sectors, a greater fraction of firms fund growth by external long-term finance.

[4] In our data, for instance, stock market returns are negatively correlated with judicial efficiency when one does not control for risk, but the relationship becomes positive when one controls for risk.

As mentioned in the introduction, the legal environment can affect the equilibrium rate of return on equity by tempering the agency problems between managers and shareholders. This can occur in two ways. First, better legal institutions may reduce the fraction of corporate resources that managers are able to divert. For instance, legal limits to managerial discretion concerning asset sales or merger agreements may curtail the scope for dilution of minority shareholders' income rights. Second, legal rules, accounting standards and courts' efficiency determine the auditing and judicial costs that shareholders incur to keep managers in line. For example, the availability of class action suits and the possibility of voting by mail reduces the cost of shareholder activism and increases its effectiveness.[5] For brevity, we shall refer to the first effect of better legal institutions as a "reduction of private benefits" and to the second as a "reduction in auditing costs".[6]

In addition, better law enforcement may benefit companies by expanding the set of contracts with suppliers and customers that can be enforced in court. Equivalently it can reduce the cost of enforcing these contracts. Ex ante, this enables each company to draw upon a wider menu of economic transactions, which should increase its profitability. This effect differs from the previous two because it does not hinge on the existence of agency problems between the company and its financiers. We refer to it as a "pure profitability effect".

The three effects are illustrated graphically in Figures 1 to 3, which refer to the case of imperfect international integration (the underlying model in Lombardo and Pagano, 2000, encompasses perfect international integration as a special case). In all three figures, the upward sloping line is the supply of equity funds to companies: the more closely integrated the country is in world capital markets, the lower is its slope. A perfectly integrated market features a flat (perfectly elastic) supply of equity. The rate of return that companies can generate for each possible level of equity funding is instead a downward sloping locus, owing to the decreasing marginal productivity of capital: we label it the profitability schedule. When managers can extract private benefits, however, companies cannot pledge their entire profits to compensate shareholders for their funds, but only a fraction of them. The rate of return after managerial diversion determines the (inverse) demand function for equity funds. The vertical distance between the profitability schedule and the demand function is the amount of private benefits per dollar invested.

[5] Managers' opportunistic behavior may also create another cost for shareholders: more volatile earnings than warranted by technology and demand conditions, and therefore additional risk. Then shareholders will require compensation for this additional risk, if they cannot easily diversify it away. In fact, they may be unable to diversify it away precisely because of the agency problem: when the costs of keeping managers in line are large, only large shareholders are willing to incur them, since they internalize the benefits of monitoring sufficiently.

[6] Lombardo and Pagano (2000) assume that expected legal and auditing costs are *proportional* to the initial investment, so that they can disregard subtle theoretical issues that arise if small shareholders free-ride on the monitoring activity of large shareholders. However, this simplifying assumption does not affect the qualitative results of their analysis. Even if small shareholders don't incur monitoring expenses, the expected costs to keep managers in check will enter equilibrium expected returns. This is because large shareholders must be compensated for the anticipated monitoring costs, and in this scenario the large investor is the *marginal* investor for pricing purposes. In addition, small shareholders may still require a premium in jurisdictions where their interests are not well protected, because they anticipate the cost of checking that large shareholders do not collude with managers. These proportional costs can alternatively be viewed as *load fees* charged by a mutual fund that shareholders join in order to overcome the free-riding problem associated with monitoring expenses. Naturally, where expected monitoring costs are higher, the mutual fund managers will charge higher fees.

The equilibrium rate of return is found at the intersection between the demand and the supply of equity funds. The cost of capital to the company (gross of private benefits) is read on the corresponding point on the profitability schedule and determines the real investment decisions of its management.

A reduction of private benefits reduces the wedge between the profitability schedule and the demand schedule. As illustrated in Figure 1, this increases both the equilibrium rate of return and the quantity of external equity, while it reduces the cost of capital to companies. The effect on the rate of return, however, is smaller the flatter is the supply schedule. In the limit, in a perfectly integrated stock market, the effect vanishes altogether (while that on the quantity is maximal).

Conversely, a reduction in the auditing costs shifts the supply curve down and to the right. As shown in Figure 2, this reduces the equilibrium rate of return, and increases the quantity of external finance. Again, the cost of capital to companies decreases. The directions of these effects are independent of the degree of equity integration: they persist even if the supply schedule is perfectly elastic.

Finally, if an improvement in the legal environment has a pure profitability effect, it will cause an outward shift of both the profitability and the demand schedules. As illustrated in Figure 3, this will increase both the observed rate of return and the amount of equity financing (the cost of equity capital may change in either direction). If the supply curve is flat, the increase in profitability does not result into an increase in the observed rate of return, as in the case of a private benefit reduction.

In summary, in an internationally integrated stock market, the effect of better institutions on the rate of return is either zero or negative. With segmented markets, their effect is predicted to be positive if they reduce private benefits or increase firm profitability, and negative if they lower the auditing costs for which investors need to be compensated. The effect on the equilibrium quantity is always positive, consistently with the evidence of LLSV (1997). However, this increase in the breadth of the equity market is not necessarily associated with a reduction in the agency costs of external finance, as shown in Figure 3.

3. Data

We are interested in capturing the empirical relationship between the legal protection of shareholders' rights and the return on equity. To this end, we use data from different sources. To measure the return on equity, we rely on three different types of data: total returns on national stock markets, accounting measures of the return on equity, and the excess return on primary equity issues as measured by IPO underpricing.

3.1. Returns on National Stock Markets

Our sample of national stock markets includes data for both developed markets from Morgan Stanley Capital International (MSCI) and emerging markets from the Emerging

Market Database (EMDB), provided by the International Finance Corporation (IFC) of the World Bank.

We draw monthly equity indices for 21 developed countries from MSCI. For eighteen of these, the sample starts on December 1969. For the remaining three, it starts on December 1987 (see Table A1 for summary statistics about MSCI markets). All indices extend to December 1999. These indices are value-weighted and are calculated with dividend reinvestment. MSCI also provides a value-weighted World Index, which serves as the market portfolio for developed countries.[7] As far as emerging markets are concerned, the starting date of coverage by EMDB differs more significantly across markets (see Table A2 for summary statistics related to the emerging markets sub-sample). For all countries, the sample extends to December 1999. As for MSCI data, the indices are value-weighted and calculated with dividend reinvestment.[8]

All returns are expressed in US dollars and are calculated in excess of the yield on the US Treasury bill that is closest to 30 days to maturity on the last trading day of the month. This latter yield is drawn from the CRSP government bond file (see Fama (1984) for the computation of holding period returns). This yield is available to us up to November 1999. As a consequence, our sample of total excess returns also ends at this date.

3.2. Accounting Measures of the Return on Capital

Data on valuation ratios (such as price-earning ratios, price-book value ratios and dividend yield) are available for the same countries from the respective above-mentioned sources on an annual basis. We also have data from the IBES global aggregates. The IBES database contains monthly valuation indices for selected (mainly developed) countries as well as survey estimates of the expected growth in earnings per share. In our estimations (see results in table 5 and 6) we use the price/earnings ratios and the dividend yields as dependent variables. The price/earnings series is defined (see Datastream International for details) as the "weighted average price/earnings ratio based on 12-month forward earnings" (IBES datatype: A12PE). The dividend yield is defined as "weighted dividend yield based on the indicated annual dividend" (IBES datatype: ADVYLD). Furthermore we obtain series on the yield on domestic 10-year government bonds (IBES datatype: AGBYLD) for a sample of 18 developed countries.

[7] The MSCI indices are broadly representative of each country's market composition. Virtually all the stocks (99%) can be traded by non-nationals as well as by domestic investors. As noted by Harvey (1991), the returns computed on the basis of these indices are highly correlated with widely quoted country indices, such as the NYSE value-weighted return (calculated by the Center for Research in Security Prices (CRSP) at the University of Chicago) for the USA, or the Nikkei 255 index for Japan. For details on the methodology behind the MSCI indexes, see MSCI.

[8] For the IFC methodology, see IFC. The selection criteria of the components of the MSCI and the IFC national indices are similar, though not identical. Bekaert and Harvey (1995) describe the EMDB indexes and briefly compare the IFC and MSCI methodologies.

3.3. The Cost of Capital on the Primary Market

We use the IPO underpricing as a proxy of the total cost of capital for firms tapping the equity capital markets for the first time. For the IPO data, our source is a study by Loughran, Ritter and Rydqvist (1994), updated by Ritter (1998). These authors provide a collection of estimates of average "IPO underpricing" in 32 countries.

3.4. Institutional Variables

Here we rely on the data set constructed by La Porta *et al.* (1998a). For a sample of 49 (both developing and developed) countries, they provide data on variables that capture: (i) the legal protection of both creditors' and minority shareholders' property rights; (ii) the origin of the national legal system; (iii) indices of the efficiency of legal enforcement; (iv) estimates of the quality of accounting systems. While we refer the reader to La Porta *et al.* (1998a) for a complete description of this interesting database and of their sources, here we briefly describe the variables that we use in our regressions and their original sources.

The variable "Judicial Efficiency" is an assessment of the "efficiency and integrity of the legal environment as it affects business, particularly foreign firms". It is produced by the country-risk rating agency Business International Corporation and is an average between 1980 and 1983, ranging from 0 to 10, with higher values associated with higher efficiency levels.

Other variables that we use were constructed by La Porta *et al.* (1998a). Among these, the index "Anti-Director Rights" captures the degree of legal protection from expropriation by the managers and controlling shareholders granted to minority shareholders. It ranges from 0 to 6, with higher scores representing more thorough legal protection of minority shareholders. The dummy variable "One Share/One Vote" equals one if in the country concerned ordinary shares are required to carry only one vote per share and 0 otherwise. The variables "French Origin", "German Origin", "Scandinavian Origin", and "English Origin" are meant to indicate the "family" to which the legal system of a given country belongs. These "legal origin" dummies may capture residual cross-country differences in investor protection on top of those specifically reflected in the "Anti-Director Rights" indicator.

Most other legal and institutional variables used here are produced by the country-risk rating agency International Country Risk (ICR). Each variable is measured as the average of the months of April and October of the corresponding ICR monthly index between 1982 and 1995, and ranges on a scale from 0 to 10 (in some cases, by re-scaling the original ICR indices). The variable "Rule of Law" is ICR's "evaluation of the legal and order tradition in the country", with lower scores for countries with weaker legal and order tradition. "Corruption" is an assessment of the degree of corruption in the government, with lower scores indicating higher corruption. The "Risk of Contract Repudiation by the Government" is an assessment of the "risk of a modification in a contract taking the form of a repudiation, postponement or scaling down" due to "budget cutbacks, indigenization pressure, a change in the government, or a change in government economic and social objectives". Lower scores for this variable indicate higher risk. The variable "Risk of Expropriation" reflects ICR's evaluation of the risk of "outright confiscation" or forced nationalization, with lower scores indicating higher risk. Since the year-by-year values of these ICR variables are available for

most of our sample periods, all the estimates reported in the next section were also repeated using these yearly values instead of their averages. The results – not reported for brevity – are qualitatively unchanged relative to those reported below.

Finally, the variable "Quality of Accounting Standards" is drawn from International Accounting and Auditing Trends (Center for International Financial Analysis & Research, Inc.), and measures the quality of companies' annual reports along 7 general dimensions (general information, income statements, balance sheets, funds flow statement, accounting standards, stock data and special items). In our complete sample of developed and emerging markets, it ranges between 24 and 83, with a mean of 61 and a standard deviation of 13.5.

4. Results

In this section we report our estimation results, divided in three subsections, each devoted to one of three alternative measures of the return on equity. In each subsection we describe the empirical methodology of our tests. In subsection 4.1, we describe our findings on the relationship between legal protection of minority shareholders and the return on equity on secondary markets. The main problem we tackle in this case is the correction for risk. In subsection 4.2, we report our results on the relationship between the institutional variables and the different accounting measures of the return on equity, such as the dividend yield and the earnings-price ratios, controlling for the effect of cross-country differences in expected earnings growth. Finally, we check if legal and judicial variables also matter for the primary equity market by investigating their correlation with estimates of IPO underpricing in different countries provided in the literature.

4.1. The Secondary Market Return on Equity

A basic tenet of asset pricing theory is that the return on any asset can be decomposed in two parts: the return on a "risk-free" asset and the compensation for undiversifiable risk. The asset pricing models proposed in the literature differ as to the quantification of the undiversifiable risk. The Capital Asset Pricing Model (CAPM) expresses it as proportional to the market beta of the asset, the proportionality factor being the market price for risk.

In the presence of agency problems between managers and shareholders, the expected returns on equity may include a third component, as explained in Section 2. While this third component may be hard to detect in any given country – where all firms operate under the same jurisdiction – it should be easier to detect and quantify in a cross-section of countries, where one can exploit the wide international variation in legal and judicial institutions which affect the severity of these agency problems.[9] Our approach will explore if such variation in

[9] Even in the context of a single country, agency costs can differ systematically across companies featuring different levels of investment in intangible assets (such as R&D). This can conceivably induce cross-sectional differences in profitability and in the cost of equity capital. For example, ceteris paribus, investors may require a large rate of return from an upstart

the respect for the law, judicial enforcement, protection of minority shareholders, and accounting standards can account for some of the international differences in the return to equity, after controlling for risk.

Indeed, such a prediction would appear to be confirmed by a quick look at Figures 4 and 5. There, we plot the average excess returns (over the period 1987-1999, chosen to maximize the number of countries included) against the efficiency of the judicial system (Figure 4) and the "rule of law" (Figure 5). In both cases, a negative *unconditional* correlation emerges from the data. However, this may simply be a reflection of the pattern of riskiness across countries: we need to purge our excess return measures of the risk premium, before we can conclude anything about the relationships we are interested in. A regression analysis is therefore required.

The basic econometric challenge in our approach is the measurement of the risk-adjusted rate of return on capital. This measurement depends crucially on the assumptions one is willing to make on the degree of integration of stock markets around the world as well as on the validity of the Purchasing Power Parity (PPP) hypothesis.

If capital markets are fully integrated internationally and PPP holds continuously, the CAPM predicts that the risk premium on each country's stock market (as measured by the excess return denominated in US dollars relative to the US risk-free rate) is proportional to that country's Beta with respect to a world market portfolio. In fact, under the PPP hypothesis, inflation differentials between countries are precisely offset by the depreciation of their bilateral exchange rate. If PPP were not to hold, then the real exchange rate risk of each country would be an additional risk factor priced on world stock markets. Adler and Dumas (1983), Harvey (1991) and Dumas and Solnik (1995), among others, show that for developed markets exchange rate risk is indeed priced. Ferson and Harvey (1993, 1994) recommend the use of a two-factor model, by adding to the traditional factor (the return on a portfolio of international stock market indices) the return on a portfolio of deposits in different currencies with weights reflecting the world trade structure.

Also the assumption of integrated capital markets may fail in practice. While there might be reasons to believe that international capital markets are increasingly integrated (owing to the removal of capital controls in developing countries and the technology-driven reduction of communication costs across borders), there is evidence that the process is gradual, not complete and not unidirectional. If equity markets are internationally segmented, a country's Beta does not (fully) capture the risk premium on its equity market, and in addition the real rate of return on the risk-free asset may differ across countries. In this case the idiosyncratic (country-specific) component of total risk should have explanatory power in the cross-section of expected returns. As mentioned in the introduction, the literature indeed shows that the "pure" international CAPM model typically has low explanatory power. Bekaert and Harvey (1995) attribute its failure to the segmentation of national markets. Harvey (1995) offers evidence on segmentation particularly for the emerging markets, by showing that the average

company with a large fraction of intangible assets, since the managers' ability to extract private benefits may be higher than in more mature companies. In the U.S., the literature has widely documented the role of firms' attributes like the ratio of market to book value or the quota of intangible assets over total value in the explanation of cross-sections of expected returns (see for example Fama and French (1993)). Some of the "anomalies" in this literature have been interpreted along the lines of agency theory.

return of these markets is positively correlated with the volatility of the market itself but not with its Beta relative to the world portfolio.

We try to take the possibility of international segmentation into account by including measures of country-specific residual risk among our regressors. We also estimate separately our regressions for the sub-samples of developed markets, on the assumption that (at least in relative terms) for emerging markets segmentation may be more of an issue.

We use two alternative methods to carry out our tests, which differ in the way we purge the cross-section of unconditional expected returns from their risk premia. The two methods are respectively in the footsteps of the time-honored multiple-steps procedures proposed by Lintner (as reproduced in Douglas (1968)) and Fama and MacBeth (1973).

The first method consists of (i) estimating for each country a time-series regression of its market's excess return on the excess return of the world market portfolio (and possibly other risk factors), and (ii) regressing cross-sectionally the unconditional averages of the N countries' excess returns (computed on the entire time interval) on the vector(s) of Beta(s) estimated in step (i), on the sample estimates of country-specific residual risk and on the institutional variables reported by La Porta *et al.* (1998a).

The second method, sometimes called "cross-sectional regression (CSR) method", involves for each month (i) estimating Betas on a preceding sample of, say, 5 years of data, (ii) running cross-sectional regressions of excess realized returns on the estimated Betas, on estimates of residual risk and – in our case – on institutional variables. The final estimated coefficient is the time average of the coefficients obtained in step (ii), taking into account the relative precision of the estimates (as explained below). Under the assumption of normality of excess returns, these "averaged" coefficients, once divided by their standard deviations, are distributed as *t*-statistics, thus allowing simple *t*-tests for inference purposes.

4.1.1. Lintner Regressions: Developed Markets

In Table 1, we report the results of the first estimation procedure, for the sample of developed markets. The sample includes all countries that are present in the MSCI database from 1970, and extends to November 1999.[10] We estimate a first-stage regression of the monthly total excess returns on the excess return of the world market portfolio and on an exchange rate risk factor. The monthly excess return of the world market portfolio is the MSCI-provided world index.[11] The exchange rate risk factor is the change in the log of the G-10 exchange rate index (obtained from the Federal Reserve System Web page). The G-10 exchange rate index is a trade-weighted average of the bilateral exchange rate between the US dollar and the 10 main trading economies (the G-7 countries, not including the US, plus the

[10] As said above, we only have data for the "Fama risk-free rate" up to November 1999.

[11] We also used the value-weighted average of the excess returns on all the markets in the MSCI and EMDB databases that are active in that month, with no qualitative change in results.

Netherlands, Belgium, Sweden and Switzerland). The change in the log of this index approximates the excess return on a trade-weighted portfolio of foreign-currency bonds, assuming that the trade weights are known and that a trade-weighted combination of foreign currency deposit rates in the 10 countries is close to the US bill rate.[12] A positive change in the G-10 index indicates a depreciation of the dollar.

This first-stage regression is used to obtain estimates for the beta and the real exchange rate risk sensitivity of each market. In column 1 of Table 1, we report our preferred specification of the international asset pricing model for developed markets. As expected, beta enters with a positive coefficient (the average market price for risk is .01 on a monthly basis). Exchange risk enters with a negative sign, as expected: market indices that are positively correlated with depreciation in the dollar exchange rate are less risky (since they offer hedging against loss in the value of the dollar) and hence offer a lower excess return.[13]

In the specification of column 2, we introduce "Judicial Efficiency" among the explanatory variables. The coefficient of this variable is positive and significant, indicating that in countries with more efficient judiciary systems the excess return on equity capital is larger. This result is clearly the opposite of what suggested by Figure 4. The estimated value (.0007) of the coefficient on this variable implies that a move from the average value for these markets of 9.37 to a perfect 10 would be associated with a 53 basis point increase of the required return on a yearly basis. It should be noticed that "Judicial Efficiency" has rather little variation among the developed markets. In most of them it is at the highest level of 10. Italy and Spain are the only countries where it is substantially lower – around 6. This implies an astonishing result for Italy and Spain: if their judicial efficiency had been in line with other countries its required return on equity would have been raised by about 3.4% per annum. As we shall see, the sign and precision of the estimated effect of "Judicial Efficiency" carries over well beyond this sample.

In column 3, we include the index "Anti-Director Rights" among the explanatory variables. Its coefficient is not significantly different from zero. In columns 4 and 5, we repeat the specifications in columns 2 and 3, after replacing the index "Judicial Efficiency" with the variable "Rule of Law", and find similar results. The point estimate for the coefficient of Rule of Law is positive, although not precisely estimated. As in column 3, the estimated coefficient of the anti-director rights variable is not statistically different from zero.

We also tried the specifications for columns 2 and 3 using other indices from La Porta *et al.* (1998a), such as measures of corruption in the government, risk of expropriation, risk of contract repudiation by the government, and the perceived quality of accounting standards. The small number of observations and the high degree of collinearity of the indices prevented us from including all of them simultaneously in our specifications. The coefficients of these

[12] Ferson and Harvey (1993) adopt a mimicking portfolio approach, *i.e.* they construct a portfolio that is maximally correlated to the change in the log of G-10, and they compute its excess return. When they use this portfolio excess return instead of the simple change in the log of G-10, they obtain similar results.

[13] We tested for deviations from the CAPM, by including among the regressors the square of beta and the variance of the residuals of our market model regressions – a proxy of unsystematic risk. The coefficients of both variables were not significantly different from zero.

indices are not significantly different from zero, while the coefficients on the two risk-factors included in the baseline specification of column 1 are virtually unchanged.

We also estimated a specification (not reported) that includes the legal origin dummies among the regressors. None of them turned out with a statistically significant coefficient, and the estimates of all the other coefficients remained qualitatively unaffected. However, we fear that this may be at least partly due to the paucity of the degrees of freedom.

4.1.2. Lintner Regressions: All Markets

In Table 2, we extend the sample size to include as many markets as possible, while requiring all markets to be included over the same time interval and at the same time keeping the interval long enough to allow reasonably precise estimates of the risk factors' sensitivities. To balance these conflicting desiderata, the best strategy appeared to be to include all the markets that enter the MSCI and EMDB databases before 1988. We compute the average excess returns and the sensitivities using monthly data over the interval between January 1987 and November 1999.[14]

Including both developed and emerging markets in the same empirical asset pricing specification presents some challenges. The empirical international finance literature, briefly surveyed above, has consistently found that possibly different risk factors are priced in the two subsets of markets. This can be seen also in Figure 6 (which plots the capital market lines for MSCI and EMDB countries) and Figure 7 (which plots the relationships between average excess returns and idiosyncratic risk in developed and emerging markets). Therefore we adapt a pragmatic approach and let the data guide us in the selection of the relevant factors needed to purge the excess returns of their risk premium component.

In the specification of column 1 of Table 2 we let the estimated values for betas, real exchange rate risk sensitivities and idiosyncratic risks free to affect differently the average returns in the two sub-samples of countries.[15] We cannot reject the hypothesis that the coefficients of the real exchange risk sensitivity, of the idiosyncratic risk for developed markets and of the betas for emerging markets are jointly equal to zero. Therefore, we adopt the constrained specification of column 2 as our baseline specification.

In column 3 we include the measure of judicial efficiency, and its estimated coefficient is positive, though not precisely estimated. In column 4 we include also the index for the legal protection of minority shareholders, and its coefficient is negative but not significant.

In column 5 and 6 we perform the same tests using the index for the rule of law as the measure of institutional quality. In both columns 5 and 6, the estimated coefficient is positive

[14] Including Finland, Ireland and New Zealand, that enter the MSCI database in 1988, does not alter the results.

[15] We also included in the list of risk factors the squares of the betas, but found that they do not have independent explanatory power. We also allowed a different intercept for EMDB countries, but its estimated coefficient was not significantly different from zero.

and significantly different from zero. In our sample for Table 2, the index for rule of law varies between 2.08 (Colombia) and a perfect ten of several developed countries (the mean is 7.26). Relative to the mean value, a one-standard deviation lower value for the index is associated with a decrease in the annualized risk-adjusted return on equity of 4.6 percentage points. In column 6, the coefficient for the index of the minority shareholders' rights is again negative but statistically indistinguishable from zero.

The pattern of results from columns 3 and 4 is found when any of the other indexes from ICRG is used. For the sake of brevity, we only report the results including the "Rule of Law" index. As a general rule, firms in countries with better values for these indices (i.e., where corruption is perceived to be less widespread or the risks of expropriation and contract repudiation are lower) reward equity capital with a higher rate of return. The accounting standards and the antidirector rights variables, instead, enter the regression with coefficients that are not significantly different from zero.

The results in both Table 1 and Table 2 provide evidence for segmentation in equity markets. As argued above, with perfectly integrated equity markets, the quality of the institutional environment is expected to affect negatively (if at all) the risk-adjusted return on equity. The model presented above helps understand through which channels better institutions can be associated with higher risk-adjusted expected returns under segmented markets. When interpreting this evidence, an important caveat should be kept in mind. Namely, the positive coefficients may result from a sample selection bias similar to that studied by Jorion and Goetzmann (1999). The true relationship between the return on equity capital and, say, judicial efficiency may be negative, at least when this variable takes very low values, but we may fail to detect it because below a minimum threshold for judicial efficiency the stock market in an emerging economy simply does not exist. In these situations, the required return on equity is effectively infinite.

4.1.3. Fama-MacBeth Cross-Sectional Regressions: Developed Markets

The empirical specifications in the above two subsections relate cross-sectional average excess returns to cross-sectional average sensitivities to the relevant risk factors and to institutional variables. A more flexible way to test whether, after controlling for the risk premium component of excess returns, a role is also played by institutional factors is to use the Generalized Least Squares methodology proposed by Litzenberger and Ramaswamy (1979), as a refinement of the Fama-Macbeth (1973) approach, widely used in explaining the cross-section of stock returns.

This procedure involves estimating an empirical model of the form:

$$r_{it} = \gamma_{0t} + \sum_{k=1}^{K} \gamma_{kt} x_{ikt} + \varepsilon_{it}, i = 1,2,...,N, t = 1,2,...,T, \quad (1)$$

where r_{it} is the excess return on country index *i* for month *t*, x_{ikt} is the k^{th} component of the vector $\vec{x}_{it}$ which includes the sensitivities of country index *i* to the relevant risk factors estimated for month *t* and (a vector of) the institutional variables for country *i* at time *t*, and ε_{it} is a disturbance with variance possibly changing across *i* and *t*. *N* is the number of

countries included in the estimation sample, and T is the number of time data points (*i.e.* the number of monthly observations used in estimation). In equation (1) the sensitivities for month t are estimated over the previous 5 years of monthly data (*i.e.* on the data for months *t-60* up to *t-1*).

The original Fama-Macbeth procedure consists of estimating equation (1) cross-sectionally for each month, so as to obtain a time series of estimated coefficients $\hat{\gamma}_{kt}$ for γ_k, k=1,..,K. If each of the estimates is assumed to be drawn from a stationary distribution, then the pooled estimates $\hat{\gamma}_k$ are:

$$\hat{\gamma}_k = \frac{1}{T}\sum_{t=1}^{T}\hat{\gamma}_{kt} \tag{2}$$

and their variance is:

$$Var(\hat{\gamma}_k) = \frac{\sum (\hat{\gamma}_{kt} - \hat{\gamma}_k)^2}{T(T-1)}. \tag{3}$$

However, this procedure does not take into account that the slope coefficients in the cross-sections for different months are estimated with different precision. The refinement proposed by Litzenberger and Ramaswamy (1979) addresses exactly this problem. They show that if the monthly estimators $\hat{\gamma}_{kt}$ are serially uncorrelated, the pooled GLS estimators $\hat{\gamma}_k$ are the weighted means of the monthly estimates, where the weights are inversely proportional to the variances of these estimates. Specifically:

$$\hat{\gamma}_k = \sum_{t=1}^{T} Z_{kt}\hat{\gamma}_{kt} \text{ where } Z_{kt} = \frac{\left[\mathrm{Var}(\hat{\gamma}_{\mathrm{kt}})\right]^{-1}}{\sum_{t=1}^{T}\left[\mathrm{Var}(\hat{\gamma}_{\mathrm{kt}})\right]^{-1}} \tag{4}$$

and

$$Var(\hat{\gamma}_k) = \sum_{t=1}^{T} Z_{kt}^2 Var(\hat{\gamma}_{kt}). \tag{5}$$

In Table 3 we report the results from this estimation procedure for the 18 developed markets for which return data are available from January 1970 in the MSCI database. Because we use 5 years of preceding data to obtain the estimates for the betas and the real exchange risk sensitivities, we run 299 cross-section regressions for the months from January 1975 to November 1999.

In column 1, we report our preferred specification for the asset pricing model. In line with column 1 of Table 1, in our baseline specification for the risk-premium we include both a country's beta relative to the excess return on the world market portfolio and a country's sensitivity with respect to real exchange rate change. We find that the market prices for these

two sources of risk have the expected sign, although the estimated market price for covariance risk is not statistically different from zero.[16,17]

In column 2 we insert also the index for the efficiency of the judicial system. As in Table 1, we find that the equity markets of countries with more efficient judicial systems pay a higher total excess return. In column 3 we include the score "Anti-Director Rights" among the regressors and find that its coefficient is very small and not significantly different from zero. Columns 4 and 5 repeat the specifications of columns 2 and 3, replacing the efficiency of the judicial system with the "Rule of Law" index. As in Table 1, the estimated coefficient of the latter variable is positive, though estimated less precisely than that of judicial efficiency.

In specifications that we do not report for brevity, we also investigate the relationship between the risk-adjusted return on equity and the degree of perceived corruption among the government officials. We find that in countries where corruption is less of a problem, the excess return over and above the reward for risk is higher, consistently with the results from the Lintner estimation procedure reported in subsection 4.1.1. The only ICRG index which is weakly negatively associated with the risk-adjusted return is the measure of the risk of expropriation. This index varies very little in the sub-sample of developed markets included in Table 3. The weakly negative coefficient appears to be sensitive to the inclusion of Hong Kong in the sample.[18]

The dummy variables for the origin of the legal systems, instead, never have any explanatory power in conjunction with the other institutional variables, particularly with the rule-of-law index, and for brevity we do not report the corresponding specifications.

4.1.4. Fama-Macbeth Cross-sectional Regressions: All Markets

In Table 4 we report results from the Lintzerberger-Ramaswamy estimation procedure for the same sample covered in Table 2, i.e. the cross section of the developed and emerging markets whose returns data are available at least from 1982.[19]

As usual, in column 1 we propose our preferred specification for the underlying model of the risk premium. In our specification search, we proceeded as for Table 2. We started by allowing beta, the real exchange risk sensitivity and the idiosyncratic risk to impact

[16] The estimated coefficients on beta may fail to pass the statistical significance test in this formulation due to inherent non-linearities in the relationship between excess returns and betas. We have considered a quadratic form for beta (along the lines of the specification used by the original Fama-Macbeth (1973) article). Although we did find some evidence for non-linearity, the results for the institutional variables reported below were qualitatively unaffected.

[17] We test and reject the hypothesis that the idiosyncratic risk has additional explanatory power, and therefore this variable is not included in the results of Table 3.

[18] Indeed, if one excludes Hong Kong from the sample, the coefficient on the risk of expropriation is positive (higher returns in countries with lower risk), and statistically significant.

[19] We want our estimation to cover the same period (1987-1999) as in Table 2. We require 5 years of data to estimate the betas and hence include a country in the sample only if we have its rate of return starting on or before 1982.

differently the expected excess return in the two sub-samples (developed and emerging markets). We found that one cannot reject the hypotheses that the idiosyncratic risk matters only for emerging markets, that the impact of beta is the same across the two sub-samples and that, controlling for these risk factors, the real exchange risk factor does not have explanatory power.

In column 2 we extend the list of regressors to include the score for "Judicial Efficiency" and, in line with the results in the previous sections, we find it positively related to the risk-adjusted excess returns. In column 3, we include the index "Anti-Director Rights" together with the measure of judicial efficiency, and find that its coefficient is small and statistically not different from zero. In column 4 and 5 we repeat the specifications of columns 2 and 3, replacing the "Judicial Efficiency" index with the score for "Rule of Law". We find the same pattern, with the index for rule of law positively associated with the risk-adjusted return on equity and, controlling for the latter, no statistically significant effects for the index of "Anti-Director Rights". Agency problem considerations – as least insofar as they are proxied by the synthetic index constructed by La Porta *et al.* (1998a) – do not appear to matter for the cross-section of excess returns.

The main message of the last two subsections is that the basic results obtained with the Lintner procedure in subsections 4.1.1 and 4.1.2 are robust to time variation in both the market price for risk and in the betas. Ferson and Harvey (1997) argue that, before concluding that a significant coefficient on a variable other than beta represents a rejection of the traditional CAPM, one needs to make sure that the variable itself has no informational content for the cross-section of betas, whose true value is, after all, unobservable. In terms of our exercise, good institutions may translate into higher expected returns because our measures of institutional quality have informational content for beta. However, our measures of institutional quality do not change over time. Hence, it is hard to ascribe their effect on expected returns to their impact on time-varying betas. We are inclined to conclude from the data that expected returns are correlated with cross-country institutional differences even after one controls for risk.[20]

4.2. Accounting Measures of the Cost of Capital

From an accounting standpoint, the profitability of a listed company is often measured by valuation ratios such as the dividend yield and the earnings/price ratio. In equilibrium, this profitability coincides with the return that shareholders require to hold the shares of the company in their portfolio. By its nature, this measure needs to be adjusted, in an international comparison, for different inflation and growth prospects. Bekaert and Harvey (1997) and Errunza and Miller (1998), among others, rely on the dividend yield as a measure of the cost of capital. As pointed out by Bekaert and Harvey (1997, p. 9), "the dividend yield has the

[20] There is no reason to expect that we underestimate the beta of countries with good institutions and overestimate it for countries with poor institutions, as would be required to interpret the positive effect of institutions on expected returns along the lines of compensation for "hidden" risk.

advantage of being directly measurable – that is, it need not be pre-estimated – and being a stationary variable."

In this sub-section, we use an augmented version of the so-called "Gordon model" of security valuation to relate the cross-country dispersion in the return on equity to international differences in legal and institutional settings. Under fairly general assumptions, the stock market index in country *i* at time *t*, P_{it}, is the expected value of discounted dividends from the component stocks into the indefinite future:

$$P_{i,t} = E_t \sum_{j=1}^{\infty} \frac{D_{i,t+j}}{(1+k_{i,t+j})^j} \tag{6}$$

where E_t is the expectation conditional on information known at time *t*, $D_{i,t+j}$ is the dividend paid out by the companies listed in country *i* at time *t+j*, and $k_{i,t+j}$ is the per-period risk-adjusted discount factor between time *t* and *t+j* relevant for the stream of dividends from country *i*. The simplest version of this valuation approach assumes k_i to be constant and $E_t(D_{i,t+j}) = D_{i,t}(1+g_i)^j$, where $D_{i,t}$ is the current dividend and g_i is a constant growth rate specific to country *i*. Under these assumptions, the expected dividend yield becomes:

$$\frac{D_{i,t}}{P_{i,t}} = k_i - g_i \,. \tag{7}$$

Of course, to compute the risk-adjusted required rate of return one needs a model of the equilibrium determination of the rates of return on stocks with different risk characteristics. Under the international CAPM, the risk of stock market index *i* is only due to its covariance with the world aggregate portfolio (and to its covariance with real exchange rate movements, if the latter is priced). If in addition the required rate of return on equities depends on the institutional variables capturing the efficiency of the legal and judiciary systems and the degree of protection of minority shareholders, k_i may be augmented as follows:

$$k_i = f_i(\beta_i) + \gamma p(L_i) \tag{8}$$

where β_i is the country's vector of betas with the relevant risk factors, $p(\cdot)$ is a polynomial and L_{it} is a vector of variables proxying for the quality of legal institutions and the degree of investor's protection in country *i*. If the classic international CAPM holds, then $f_i(\beta_i) = \alpha r_i + \eta\beta_i$, where η is the market price for risk (defined as the excess return on the risk-free rate of the world stock market), r_i is the yield on a long-term "risk-free" domestic security (in the estimation we use the yield on a 10-year government bond, drawn from the IBES database), and $\gamma = 0$. The nominal domestic interest rate must be included, since our estimates of the expected growth in earnings per share are in nominal terms and the dividend yields are denominated in different currencies. As a result, we must allow for different yields to reflect different expected inflation rates, even if all other factors were the same across markets. In the estimation, we shall allow for time-varying k_i's, that is, we shall let the expected return on country *i*'s stock market index vary over time. However, the fact that for some dates data are available only for a few markets prevents a cross-sectional regression

approach and therefore limits the flexibility of the functional form for the required rate of return in equation (8). We impose that the market price for risk and the coefficient γ on the "institutional variables" be constant, that is:

$$k_{i,t} = \alpha r_{it} + \eta \beta_{it} + \gamma p(L_{it})$$

Imposing this linear specification for the risk premium component of the return on equity in equation (8), one obtains the specification to be estimated:

$$\frac{D_{i,t}}{P_{i,t}} = \alpha r_{it} + \eta \beta_{it} + \gamma p(L_{it}) + \delta g_i + \varepsilon_{it}, \qquad (9)$$

where the restriction $\delta = -1$ should hold if (i) this simple version of Gordon's model were true and (ii) if our measure of dividend growth were free of measurement errors. In fact, we expect neither of these to be true. First, the expected growth of dividends is unlikely to be constant, as assumed in the derivation of (7). In addition, our proxy of expected dividend growth – being a survey-based measure – may be vitiated by measurement errors, due for example to the selection of survey respondents, or to imperfect coincidence between their reference portfolio and the country portfolio used to construct our dividend yields. Last but not least, we proxy dividend growth with earnings growth, which is inappropriate unless payout ratios are constant. For all these reasons, we do not expect the restriction $\delta = -1$ to hold in our regressions.

Equation (9) can be estimated for each time period for which we have data on the dividend yield of the stock market index. It requires an estimate for β_{it}. This, as before, can be obtained by a first-stage regression of the market *i* total (excess) return on the world excess return up to a time period strictly before *t*, so as to avoid covariance between β_{it} and ε_{it}. In the empirical specification, we allow for the possibility of cross-market correlation and heteroskedasticity for the errors ε_{it}, as well as for clustering of the error terms within each country.

Notice also that from equation (9) different specifications can be derived, which involve other familiar (and widely used in practice) accounting measures of value. For example, if one assumes that dividends are a common fraction of earnings in all countries, then one immediately obtains another testable specification involving earning-price ratios:

$$\frac{E_{i,t}}{P_{i,t}} = \alpha' r_{it} + \eta' \beta_{it} + \gamma'_t \, p(L_{it}) + \delta' g_i + \varepsilon'_{it}, \qquad (10)$$

where the superscript primes indicate that the coefficients in (10) may differ from the analogous coefficients in (9) because of division by the payout ratio *b*, assumed to be common across countries and constant over time. Since, however, differences in tax treatment of dividends across countries may affect payout ratios, we include a measure of the relative stance of the tax system towards different uses of earnings as a separate regressor in our estimation. The measure is drawn from La Porta *et al.* (1998b), who actually find that payout ratios are only tenuously correlated with the tax disadvantage of dividends relative to capital gains.

4.2.1. Dividend Yields

Our empirical results for the dividend yields are reported in Table 5. Here we estimate equation (9) with monthly observations on the dividend yields for 18 developed markets. We cannot include emerging markets for lack of data. In the columns of this table we report various specifications, which include different institutional variables in the vector *p(L)*. To interpret our results, it is useful to keep in mind the scale of the dependent variable in our regression. In our sample, the average dividend yield is 0.027, its standard deviation is 0.0115 and its range is between 0.002 and 0.079.

In column 1, we report our baseline specification. We include the domestic government bond yield, the expected growth in earnings per share and the market beta as the determinants of the required rate of return. We estimate market *i*'s beta for month *t* by regressing market *i*'s excess return on the world market portfolio's excess return for the previous 60 months. Consistently with our theoretical model, we find positive coefficients on the domestic government bond yield, and a negative (though imprecisely estimated) coefficient for the expected growth in earnings per share. These results hold in all the specifications of Table 5.[21] In our baseline specification, we also include the legal origin dummies, and the dummy variable "One Share/One Vote". We allow for the differential tax treatment of dividends to have an impact on the dividend policy of firms, by including the variable "Dividend Tax Preference" from La Porta *et al.* (1998b). This variable is defined as the ratio of the net-of-taxes value to outside shareholders of 1 dollar in earnings distributed out as dividends to the net-of-taxes value of 1 dollar of earnings retained in the firm. It is meant to capture the extent of tax disadvantage borne by dividends relative to capital gains.

Even after controlling for undiversifiable risk and adjusting the accounting measures for differences in expected inflation and growth prospects, the legal origin dummies play an important role in explaining the cross-sectional behavior of the dividend yield. Using this measure of the return on equity, countries with German and Scandinavian legal systems have lower risk-adjusted equity returns than English-origin countries. The coefficient of the "One Share/One Vote" variable is negative, consistently with agency theories of corporate governance. The coefficient of the tax variable has the expected positive sign, although it is not precisely estimated (and it is small: a move from the average 0.75 to a non-distorting value of 1 would increase the dividend yield by 0.0001, or 8.4% of a standard deviation).[22] The regression has a good fit: the R^2 for the specification in column 1 is 0.635.

In column 2, we further include the index for the degree of legal protection of minority shareholders from managers' opportunism ("Anti-Director Rights"). Consistently with our

[21] The estimated market price for risk is positive, although not always statistically significant.

[22] For an analysis of the possible reasons why dividend payouts across markets appear to be so little sensitive to tax considerations, see La Porta *et al.* (1998b). On an intuitive level, signalling theories of dividends (and, in particular, "burning money" theories of dividends) argue that firms use dividends exactly because they are a relatively costly way of disbursing cash out to shareholders, to signal their financial strength. Hence they might explain why in countries where dividends are more costly they are used relatively more often. See Bernheim (1990) for an example of such theories.

results for this variable in the previous section, we find a negative but statistically insignificant coefficient (-0.0013). We find that the effects of differences in the four legal origins are not fully absorbed by differences in the legal protection of minority shareholders. The coefficients of the Scandinavian and German origin dummies maintain their sign and their significance increases somewhat upon controlling for the degree of minority shareholders' protection. The coefficient on the French legal origin dummy turns negative, possibly reflecting this variable's negative correlation with the LLSV measure of minority shareholders' legal protection.[23]

In column 3, we include our measure of judicial efficiency. Consistently with our results in the previous section, we find that firms in countries with a higher degree of judicial efficiency pay a higher risk-adjusted return on their equity capital (our index has an average of 9.27, with a range between 6.25 and 10). The other variables' coefficients are significantly different from zero. In column 4 we include the anti-director rights index, and we obtain a negative estimate, once again statistically undistinguishable from zero. The R^2 in the specifications of columns 3 and 4 is a remarkable 0.71.

In columns 5 and 6 we repeat the specifications of columns 3 and 4 respectively, controlling for the rule of law index *in lieu* of the efficiency of the judicial system. Again, we find that such a variable enters with a positive and economically relevant positive coefficient.[24] At the same time, the anti-director rights index enters with a negative sign. In both columns 5 and 6 however, the coefficients on these two variables are not precisely estimated.

In column 7 we repeat the specification of column 6 (including the anti-director rights score) but we also control for the quality of accounting standards. This turns out to enter with a positive and significant coefficient.[25] The same is true for the other measures of quality of the business environment: corruption in government, risk of contract repudiation and risk of expropriation. However, the estimated coefficients are not statistically significant, which may reflect the fact that these variables exhibit limited variation across developed markets. The specifications including these latter three variables are not reported for brevity.[26]

[23] We find that including the anti-director rights score reduces the coefficient on the French dummy across all our proposed specifications (see below).

[24] The rule of law index averages 9.34 in our sample and has a range from 7.8 to a perfect 10. A move from 7.8 to 10 is associated, through our point estimate, to an increase in the dividend yield of 0.0037, or 31% of its standard deviation.

[25] The index for accounting standards averages 68.23 in our sample, and has a range between 54 and 83. Through our point estimate (0.0003) therefore this variable can potentially explain an increase in the dividend yield of 0.0003*(83-54)=0.0087, or 75% of its standard deviation.

[26] We have re-estimated all the specifications included in Table 5 controlling for real exchange rate risk. The results are unaffected and are not reported for brevity.

4.2.2. Earnings-Price Ratios

In Table 6 we report our results for the earnings-price (EP) ratios. Again, data limitations constrain us to the set of developed markets. The dependent variable is defined as the reciprocal of the PE ratio as obtained from IBES.[27] From equation (10) we expect to find that the domestic government bond yield and the beta enter with a positive sign, while the expected growth rate in earnings per share with a negative one. Table 6 confirms the results from the previous subsection on the estimated impact of measured institutional differences on the international dispersion of return on equity.

In column 1, as usual, we report our baseline specification. This includes the domestic government bond yield, the expected growth in earnings per share and the beta with the world market portfolio (estimated on the previous 60 months, market by market), as well as the one-share/one-vote dummy and the origin dummies.[28] To interpret our results, the reader should keep in mind that our dependent variable averages 0.069 in the sample, has a standard deviation of 0.0216 and a range between 0.014 and 0.135. The domestic bond yield, the expected growth rate and beta all enter with coefficients whose signs are in line with our a-priori expectations (although the estimated market price for risk is again not statistically significant).

As in all the previous tables referring to developed countries, all our institutional variables enter with positive signs. The "general environment" variables (i.e. the efficiency of the judicial system, the rule of law, the degree of corruption among government officials, the risk of contract repudiation and of expropriation, the quality of accounting standards), all increase the required return on equity, *ceteris paribus*. The degree of protection of minority shareholders enters with a negative sign, but it is never precisely estimated. And, again as in Table 5, the Scandinavian and German origin dummies' impact goes in the direction of making the average return on equity lower in the corresponding countries than in the English origin one. Anglo-Saxon markets appear to have conditionally larger returns on equity than all other developed countries' markets, after controlling for both risk and measured institutional differences.

The one-share/one-vote dummy enters with a large and negative coefficient, in accordance with the predictions of the agency theory of the cost of external funds, as well as with our results in the previous subsection. For instance, Japan and Singapore, who have the one-share/one-vote compulsory requirement in their commercial codes, *ceteris paribus* have a significantly lower EP ratio than markets that lack this legal provision.

[27] We use the earnings/price ratio instead of the price/earning ratio to avoid the problems that arise when earnings are very small (see also Ferson and Harvey, 1997).

[28] In the specifications we report, we do not include the dividend preference variable, as the theory suggests that it should not matter. However, we have run the same regressions that we report in table 6 including the dividend tax preference variable. Although usually precisely estimated, its coefficient changes sign across specifications, making us somewhat skeptical as to its interpretation. It enters with a positive sign if one controls for the anti-director rights score, otherwise it enters with a negative sign.

4.3. The Excess Return on Equity in the Primary Market

In the primary market, the excess return earned on new issues in the first days immediately after the quotation coincides with the so-called initial public offering (IPO) underpricing. So in this section we use the estimates of IPO underpricing described in subsection 3.3 to investigate whether international differences in this variable can be explained by cross-country differences in the institutional variables analyzed throughout the paper. The evidence is to be taken with a grain of salt due to the paucity of the sample and the heterogeneity of the measures of IPO underpricing.

Theory suggests that differences in accounting standards should be a key explanatory variable of the international variation in IPO underpricing. The presence of IPO underpricing is generally viewed as the product of informational asymmetries between the generality of investors (the "uninformed" bidders) and the "smart money" in the market for new issues. Shares initially quote at a discount to compensate the uninformed investors for their expected losses to the better-informed ones. This informational asymmetry and the resulting IPO discount are likely to be greater where accounting practices are lax and opaque.

Figure 8 is consistent with the prediction of the theory: there is a simple negative correlation between IPO underpricing and accounting standards.[29] This result is confirmed by the regression reported in the first column of Table 7: the correlation is indeed negative and precisely estimated. In the rest of Table 7 we investigate if this result is robust to the introduction of other institutional variables among the explanatory variables. The correlation remains negative and significant when one controls for most of the other institutional variables, such as anti-director rights (whose coefficient is – surprisingly – positive, though not precisely estimated) and rule of law (column 4). In specifications that include corruption, risk of contract repudiation and the "one-share/one-vote" variable (not reported), the latter does not enter with a significant coefficient.

The magnitude and precision of the coefficient of the accounting standards variable are considerably reduced only when the measure of judicial efficiency is entered as an additional explanatory variable, presumably due to their collinearity. In fact, Figure 9 confirms a strong negative correlation between IPO underpricing and judicial efficiency. This probably captures simply the fact that where courts can be trusted to do their job honestly and efficiently, accountants can be trusted to do the same: their honesty and accuracy may partly result from the threat of swift judicial suits if they misbehave.

Altogether, the main result in this table appears to be the fact that good accounting standards appear to reduce the cost of capital on primary equity markets. The special relevance of this institutional variable in the context of the primary equity market accords well with the theory.

[29] In figures 8 and 9, some countries appear more than once. This is because, for those countries, we have data on the average IPO underpricing for different issuing procedures. The letter in parenthesis after the country's name indicates to which issuing procedures the observation refers, following the same convention as in Table 7. The letter "a (b)" indicates that in the IPO the offer price is fixed before (after) the acquisition of information and the allocation is discretionary. The letter "c" indicates that in the IPO the offer price is fixed before the acquisition of information and the allocation is not discretionary. The letter "e" signals the presence of binding regulatory constraints (see Table 2 in Loughran et al., 1994).

5. Concluding remarks

In this paper we estimate the correlation between the quality of the institutional environment and the required return on equity for a cross-section of national stock market indices from both developed and emerging markets. We use several measures of the return on equity, such as the total return on national equity markets (controlling for risk premia), accounting measures of firms' profitability such as dividend yields and earning-price ratios (controlling for international differences in growth and inflation), and IPO average excess returns (IPO underpricing).

We find an interesting set of regularities. First, when we use secondary market returns on equity, all our estimates reveal a positive correlation between the risk-adjusted return on equity and "general" measures of the quality of institutions, like efficiency of the judicial system, respect for the law, lack of corruption among government officials, quality of accounting standards and low risk of contract repudiation and nationalization. The same results are found when we use accounting measures of the rate of return on equity: both the dividend yield and the earnings-price ratio are positively correlated with these general measures of the quality of institutions.

As explained in section 2, these findings can be rationalized only in the context of imperfectly integrated equity markets. When the world capital markets are not fully integrated, the supply of funds at the national level is upward sloping, rather than perfectly horizontal. In this scenario, our findings can result from two types of effects of better institutions. First, more effective courts and higher respect for the law can reduce the amount of private benefits that the management can extract from corporate resources, thus allowing companies to credibly pledge higher returns to investors. Second, better institutions can have a positive effect on the profitability of companies, quite apart from their impact on financial relationships. Since a wider menu of contracts can be effectively enforced, the marginal productivity of physical investment is increased: there is more demand for equity funding in equilibrium and hence a higher rate of return. Notice that both types of effects would result in an increase in the amount of equity funding used by companies, consistently with the evidence in La Porta *et al.* (1997).

Our finding may appear harder to reconcile with one of the results of Demirgüc-Kunt and Maksimovic (1998): they report that the return on assets of the firms in their sample is negatively correlated with the same rule-of-law indicator used in our study, controlling for various macroeconomic variables (inflation, deposit bank assets and stock market capitalization divided by GDP, etc.). The contradiction between the two studies – apart from the different measures of profitability and data used – may however be due to the fact that Demirgüc-Kunt and Maksimovic fail to control for risk in their regression. As we know from Figure 2, the unconditional correlation between the return on equities and rule-of-law is clearly negative. Nevertheless our regressions show that, once risk is controlled for, the partial correlation between these variables is positive.

Our second key result concerns the protection of shareholder rights. We find that the return on equity is basically uncorrelated with the measure of shareholder rights proposed in La Porta et al. (1998). This may either be due to the fact that the index constructed by La Porta et al. does not fully capture all the features of the legal codes and of the regulatory

apparatus that determine the degree of protection of shareholders' rights (as argued for example by Coffee (1999)). Or it may be due to the fact that, when the law does not protect minority shareholders, large shareholders typically arise to overcome this problem (as argued by La Porta et al. (1998a)). This result parallels the one by Bhattacharya and Daouk (1999), who investigate the effect on the cost of equity capital of insider trading regulation and find that, while the mere existence of laws prohibiting insider trading is ineffectual, their enforcement reduces the risk-adjusted expected return on equity.

Thirdly, we have found that the origin of the legal code is an important determinant of the international variation in the risk-adjusted return on equity. This is an important result, because the origin dummies are, among all the variables we use, probably those that can be more safely thought to be truly exogenous. La Porta *et al.* (1997) found that countries in the English legal tradition area have a bigger stock market capitalization relative to their GDP and more external equity financing than countries in the French, German or Scandinavian legal families. We find that - if one controls for general institutional features, such as the efficiency of the judicial system, the rule of law or the quality of accounting standards - firms in the German and the Scandinavian families pay a lower risk-adjusted return than firms in Anglo-Saxon countries. If, as maintained by La Porta et al., shareholders are better protected in Anglo-Saxon countries than they are in German and Scandinavian countries, then this finding represents another piece of evidence for segmentation of international equity markets. It points to the existence of non-zero risk-adjusted differentials in expected returns, even after controlling for differences in the expected monitoring, auditing and other private costs associated with different legal environments. Companies in Anglo-Saxon countries may be able to pay a higher return because they are more profitable than comparable firms in other countries, all else being equal, including measurable institutional differences. A possible interpretation is that the "English origin" dummy captures some unmeasured characteristics in the social organization of these countries that makes their companies more profitable than companies located elsewhere. This result is of interest also because of its connection with international differences in the equity premium puzzle documented by Jorion and Goetzmann (1999).

Our fourth, and final, result concerns the cost of capital in the primary market. We find that IPO underpricing – an important component of the total cost of capital for newly listed companies – is negatively and significantly correlated with the quality of accounting standards across countries. This accords with theories of IPO underpricing based upon adverse selection, as in countries where accounting information is less reliable the informational gap between informed and uninformed investors is bound to be larger at the IPO stage.

Our analysis leaves a number of issues open to further inquiry. First, an intriguing issue is which precise mechanism generates the positive correlation between the efficiency of judicial enforcement of contracts and the return on equity. A second task is to test the robustness of our results on more disaggregated data. Lombardo (2000) estimates the effect on the rate of return on equity of changes in the institutional environment using a panel industry-level data, controlling for other country-specific sources of international differentials in the risk-adjusted expected return. He finds that, within a given country, improvements in the institutional framework significantly reduce the risk-adjusted return on equity. Third, a fascinating line of research involves endogenizing the evolution of the institutional framework, so as to understand why some countries end up having and retaining "bad" institutions (for a step in this direction, see Pagano and Volpin, 1999).

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TABLE 1

AVERAGE EXCESS RETURN, RISK AND INSTITUTIONAL VARIABLES IN DEVELOPED MARKETS

The dependent variable is the average monthly total excess return in US dollars for the markets of the MSCI database present from January 1970. The safe rate of return is the yield on the 30-day maturity government bond with closest maturity (source: "Fama files" from CRSP database). The average excess return is computed over the period from Jan. 1970 to Nov.. 1999. Beta and the Exchange Rate Factor are computed from OLS regressions estimated over the same intervals. Non-Systematic Risk is the standard deviation of the residuals obtained from the same regressions. All other variables are from La Porta et al. (1998a). T-statistics, computed with heteroskedasticity-consistent standard errors, are reported underneath the estimated coefficients.

Regressors	**1**	**2**	**3**	**4**	**5**
Beta	.010 (2.36)	.009 (2.22)	.008 (2.42)	.010 (2.79)	.009 (3.14)
Exch. Rate Factor	-.004 (-2.19)	-.0039 (-2.21)	-.004 (-1.729)	-.004 (-2.47)	-.004 (-2.18)
Judicial Efficiency	-	.0007 (2.37)	.0007 (2.03)	-	-
Rule of Law	-	-	-	.0002 (0.26)	.0001 (0.22)
Anti-director Rights	-	-	.00001 (.32)	-	.0004 (.75)
Intercept	.002 (0.66)	-.003 (-.72)	-.003 (-.73)	-.0002 (-0.02)	-.0001 (-0.01)
Number of Obs.	18	18	18	18	18
R Squared	.34	.44	.44	.34	.37

TABLE 2

AVERAGE EXCESS RETURN, RISK AND INSTITUTIONAL VARIABLES IN DEVELOPED AND EMERGING MARKETS

The dependent variable is the average monthly total excess return in US dollars for the (developed) markets of the MSCI database and the (emerging) markets of the EMDB database, available from January 1987. The safe rate of return is the yield on the 30-day maturity government bond with closest maturity (source: "Fama files" from CRSP database). The average excess return is computed over the period from Jan. 1987 to Nov. 1999. Beta and the Exchange Rate Factor are computed from OLS regressions estimated over the same intervals. Non-Systematic Risk is the standard deviation of the residuals obtained from the same regressions. These variables are interacted with dummy variables to allow for different slope coefficients in the MSCI and EMDB sub-samples. All other variables are from La Porta et al. (1998a). T-statistics, computed with heteroskedasticity-consistent standard errors, are reported underneath the estimated coefficients.

Regressors	**1**	**2**	**3**	**4**	**5**	**6**
Beta in MSCI countries	.010 (2.09)	.011 (3.59)	.010 (3.33)	.011 (3.09)	.009 (2.67)	.009 (2.45)
Exch. Rate Factor in MSCI countries	.001 (.26)	-	-	-	-	-
Non-Systematic Risk In MSCI countries	-.002 (-.04)	-	-	-	-	-
Beta in EMDB countries	-.001 (-.32)	-	-	-	-	-
Exch. Rate Factor in EMDB countries	.002 (.89)	-	-	-	-	-
Non-Systematic Risk In EMDB countries	.140 (5.847)	.147 (7.29)	.157 (7.25)	.161 (6.39)	.154 (8.22)	.156 (7.45)
Judicial Efficiency	-	-	.0005 (1.12)	.00061 (1.26)	-	-
Rule of Law	-	-	-	-	.0008 (2.185)	.0008 (2.09)
Anti-director Rights	-	-	-	-.0003 (-.60)	-	-.0002 (-.32)
Intercept	.001 (.47)	.001 (.35)	-.003 (-.76)	-.003 (-.71)	-.004 (-1.29)	-.004 (-1.07)
Number of Obs.	37	37	37	37	37	37
R Squared	.66	.65	.66	.66.	.68	.68

TABLE 3

AVERAGE SLOPES OF MONTHLY CROSS-SECTIONAL REGRESSIONS OF RETURNS ON BETA, EXCHANGE RATE RISK AND INSTITUTIONAL VARIABLES IN DEVELOPED MARKETS

Monthly total excess returns in US dollars for the markets of the MSCI database are regressed each month on the explanatory variables for the period from January 1975 to November 1997. For each month, both Beta and the exchange rate risk sensitivity are estimated on the previous 60-month period. The safe rate of return is the yield on the 30-day maturity government bond with closest maturity (source: "Fama files" from CRSP database). All other variables are from La Porta et al. (1998a). Coefficients and standard errors are computed with the pooled GLS estimators. T-statistics are reported underneath the estimated coefficients.

Regressors	**1**	**2**	**3**	**4**	**5**
Beta	.001 (.80)	.0001 (.09)	.0007 (.38)	.003 (1.67)	.003 (1.84)
Exchange Rate Risk	-.001 (-1.50)	-.002 (-1.81)	-.003 (-2.12)	-.002 (-1.65)	-.003 (-2.29)
Judicial Efficiency	-	.001 (2.16)	.001 (2.07)	-	-
Rule of Law	-	-	-	.0007 (0.94)	.0007 (.91)
Anti-director Rights	-	-	.0001 (.30)	-	.0002 (.54)
Number of Obs. in Cross-Sectional Regressions	18	18	18	18	18

TABLE 4

AVERAGE SLOPES OF MONTHLY CROSS-SECTIONAL REGRESSIONS OF RETURNS ON BETA, RESIDUAL RISK AND INSTITUTIONAL VARIABLES IN DEVELOPED AND EMERGING MARKETS

Monthly total excess returns in US dollars for equity markets present in the MSCI and EMDB databases from January 1982 are regressed each month on the explanatory variables for the period from January 1987 to November 1999. The safe rate of return is the yield on the 30-day maturity government bond with closest maturity (source: "Fama files" from CRSP database). For each year, Beta is estimated on the previous 60-month period. Non-Systematic Risk in Emerging Markets is the standard deviation of the residuals from these first-stage regressions, interacted with a dummy which equals one if the market is from the EMDB and zero otherwise. All other variables are from La Porta et al. (1998a). Coefficients and standard errors are computed with pooled GLS estimators. T-statistics are reported underneath the estimated coefficients.

Regressors	**1**	**2**	**3**	**4**	**5**
Beta	.004 (1.87)	.004 (2.17)	.004 (1.98)	.002 (2.18)	.002 (.91)
Non-Systematic Risk In Emerging Markets	.029 (1.44)	.087 (3.29)	.079 (2.96)	.101 (3.59)	.091 (3.24)
Judicial Efficiency	-	.0015 (2.49)	.0013 (2.06)	-	-
Rule of Law	-	-	-	.0015 (2.38)	.0013 (2.05)
Anti-director Rights	-	-	.0003 (.56)	-	.0004 (.83)
Number of Obs. in Cross-Sectional Regressions	28	28	28	28	28

TABLE 5

DIVIDEND YIELDS, RISK AND INSTITUTIONAL VARIABLES IN DEVELOPED MARKETS

Monthly dividend yields, expected growth in earnings per share, and government bond yields are from the IBES global aggregates database (Datastream International). The dividend yield is the weighted yield based on the indicated annual dividend (IBES datatype: ADVYLD). "Growth in Earnings Per Share" is the weighted 12-month-forward growth in earnings per share (EPS) (IBES datatype: A12GRO). "Government Bond Yield" is a generic yield based on a local 10-year government bond (IBES datatype: AGBYLD). Beta for market *i* in month *t* is estimated from market model regressions of market *i*'s excess return on the world market excess return in months *t*-1,.., *t*-60. Institutional variables are from La Porta et al. (1998a). The data are an unbalanced panel of the MSCI markets from January 1987 to November 1999. Heteroschedasticity-consistent t-statistics are reported underneath estimated coefficients.

Regressors	1	2	3	4	5	6	7
Government Bond Yield	.171 (3.77)	.177 (3.75)	.1855 (5.31)	.1877 (5.15)	.1680 (3.85)	.1740 (3.78)	.1821 (4.06)
Expected Growth in Earnings Per Share	-.0001 (-.99)	-.0001 (-.90)	-.0001 (-1.06)	-.0001 (-1.00)	-.0001 (-.983)	-.0001 (-.89)	-.0001 (-.91)
Beta	.0037 (1.01)	.0050 (1.55)	.007 (1.90)	.0075 (2.25)	.0050 (1.28)	.0061 (1.836)	.0038 (1.09)
Judicial Efficiency	-	-	.0046 (3.34)	.0045 (3.00)	-	-	-
Rule of Law	-	-	-	-	.0010 (0.518)	.0009 (0.46)	-
Quality of accounting standards	-	-	-	-	-	-	.0003 (1.92)
Anti-director Rights	-	-.0013 (-.72)	-	-.0005 (-.35)	-	-.0012 (-.71)	-.0018 (-1.05)
One Share/One Vote	-.0132 (-4.09)	-.0128 (-3.73)	-.0167 (-4.32)	-.0164 (-4.06)	-.0130 (-3.72)	-.0126 (-3.48)	-.0133 (-3.92)
French Origin	.0007 (.16)	-.003 (-.81)	.0089 (4.044)	.0072 (1.50)	.0012 (.311)	-.0023 (-0.68)	-.0019 (-.42)
German Origin	-.0073 (-2.17)	-.0106 (-2.42)	-.0052 (-1.59)	-.0066 (-1.52)	-.0074 (-2.07)	-.0106 (-2.538)	-.0080 (-1.47)
Scandinavian Origin	-.0163 (-6.82)	-.0181 (-5.79)	-.0186 (-5.65)	-.0194 (-5.73)	-.0170 (-5.79)	-.0187 (-5.92)	-.0183 (-4.34)
Dividend Tax Preference	.0038 (.47)	-.0007 (-.101)	.0139 (2.22)	.0119 (1.54)	.0048 (.53)	.0004 (.06)	-.0048 (-.694)
Intercept	.013 (1.679)	.0210 (2.13)	-.0438 (-2.60)	.0235 (-1.689)	-.0014 (-.06)	.0100 (.412)	.0080 (2.013)
Number Of Obs.	1455	1455	1455	1455	1455	1455	1394
R-squared	.635	.640	.710	.711	.638	.643	.665

TABLE 6

EARNINGS/PRICE RATIOS, RISK AND INSTITUTIONAL VARIABLES IN DEVELOPED MARKETS

Monthly data for earnings/price (EP) ratios, expected growth in earnings per share and government bond yield are from the IBES global aggregates database (Datastream International). The EP ratio is the reciprocal of the price/earnings (PE) data from IBES, defined as "Weighted average price/earnings ratio based on 12-month forward earnings" (IBES datatype: A12PE). "Expected Growth in Earnings Per Share" is the weighted 12-month-forward growth in earnings per share (EPS) (IBES datatype: A12GRO). "Government Bond Yield" is a generic yield based on a local 10-year government bond (IBES datatype: AGBYLD). Beta for market *i* in month *t* is estimated from market model regressions of market *i*'s excess return on the world market excess return in months *t*-1,.., *t*-60. Institutional variables are from La Porta et al. (1998a). The data are an unbalanced panel of the MSCI markets from March 1987 to November 1996. Heteroschedasticity-consistent t-statistics are reported underneath estimated coefficients.

Regressors	1	2	3	4	5	6	7
Government Bond Yield.	.404 (3.66)	.403 (3.69)	.4382 (4.19)	.4371 (4.26)	.401 (3.76)	.4001 (3.78)	.4239 (3.89)
Expected Growth in Earnings Per Share	-.001 (-2.46)	-.001 (-2.46)	-.001 (-2.41)	-.0009 (-2.67)	-.0010 (-2.43)	-.0010 (-2.42)	-.0010 (-2.48)
Beta	.0023 (.36)	.0028 (.42)	.0061 (.87)	.0072 (1.04)	.0024 (.83)	.0061 (.987)	-.0047 (-.66)
Judicial Efficiency	-	-	.0047 (2.54)	.0048 (2.57)	-	-	-
Rule of Law	-	-	-	-	.0024 (.77)	.0025 (.83)	-
Quality of accounting standards	-	-	-	-	-	-	.0003 (1.02)
Anti-director Rights	-	-.0006 (-.28)	-	-.0011 (-.74)	-	-.0007 (-.39)	-.0001 (-.004)
One Share/One Vote	-.0228 (-2.12)	-.0228 (-2.15)	-.0256 (2.15)	-.0255 (-2.21)	-.0222 (-1.97)	-.0220 (-2.002)	-.0202 (-1.96)
French Origin	-.0033 (-.67)	-.0046 (-.65)	.0030 (.69)	.0006 (.10)	-.0024 (-.49)	-.0041 (-.56)	.0021 (.27)
German Origin	-.0136 (-2.29)	-.0150 (-1.82)	-.0123 (-1.89)	-.0149 (-1.78)	-.0142 (-2.47)	-.0159 (-2.06)	-.0071 (-.67)
Scandinavian Origin	-.0061 (-2.46)	-.0070 (-1.579)	-.0084 (-2.90)	-.0101 (-2.698)	-.0078 (-3.59)	-.0090 (-2.88)	-.0028 (-.62)
Intercept	.0462 (4.12)	.0486 (3.46)	-.005 (-.21)	-.0023 (-.08)	.0203 (-.57)	.0225 (.21)	.0212 (0.88)
Number of Obs.	1455	1455	1455	1455	1455	1455	1394
R-squared	.620	.620	.645	.647	.624	.625	.626

TABLE 7

IPO UNDERPRICING AND INSTITUTIONAL VARIABLES

The dependent variable is the average IPO underpricing reported by Loughran, Ritter and Rydqvist (1994), as updated by Ritter (1998). "One Share/One Vote" is a dummy variable, equal to 1 if the legal system of the country explicitly imposes that each share be given one and only one vote in the shareholders' meetings, and 0 otherwise. The "dummy a (b)" equals 1 only if in the IPO the offer price is fixed before (after) the acquisition of information and the allocation is discretionary. The "dummy c" equals 1 if in the IPO the offer price is fixed before the acquisition of information and the allocation is not discretionary. The "dummy e" equals 1 only in the presence of binding regulatory constraints (see Table 2 in Loughran *et al*., 1994). Heteroschedasticity-consistent t-statistics are reported underneath estimated coefficients.

Regressors	**1**	**2**	**3**	**4**
Judicial Efficiency	-	-	-.050 (-2.531)	-
Rule of Law	-	-	-	-.033 (-1.035)
Quality of Accounting Standards	-.014 (-2.510)	-.016 (-2.704)	-.010 (-1.323)	-.015 (-1.958)
Anti-director Rights	-	.055 (1.658)	.054 (1.640)	.049 (1.454)
Dummy a	.351 (3.654)	.385 (4.155)	.362 (3.690)	.363 (3.670)
Dummy b	.117 (1.354)	.055 (0.562)	.053 (0.534)	.042 (0.425)
Dummy c	.374 (2.997)	.342 (2.771)	.290 (2.135)	.316 (2.277)
Dummy e	.962 (12.408)	.970 (9.451)	.870 (8.249)	.852 (5.351)
Intercept	.980 (12.408)	.987 (2.688)	1.048 (2.707)	1.210 (3.722)
Number of Obs.	31	31	31	31
R-squared	.650	.686	.729	.699

TABLE A1

SUMMARY STATISTICS: ANNUALIZED TOTAL RETURN IN U.S. DOLLARS

(MSCI DATABASE)

Country	Start	Arithmetic Mean	Geometric Mean	Standard Deviation
Australia	Jan, 1970	0.116	0.084	0.255
Austria	Jan, 1970	0.126	0.093	0.209
Belgium	Jan, 1970	0.162	0.145	0.188
Canada	Jan, 1970	0.114	0.098	0.185
Denmark	Jan, 1970	0.156	0.135	0.187
Finland	Jan, 1988	0.123	0.201	0.269
France	Jan, 1970	0.144	0.125	0.233
Germany	Jan, 1970	0.138	0.121	0.203
Hong Kong	Jan, 1970	0.254	0.176	0.394
Ireland	Jan, 1988	0.158	0.120	0.197
Italy	Jan, 1970	0.102	0.071	0.264
Japan	Jan, 1970	0.152	0.134	0.229
Netherlands	Jan, 1970	0.172	0.154	0.177
New Zealand	Jan, 1988	0.085	0.037	0.230
Norway	Jan, 1970	0.160	0.109	0.267
Singapore	Jan, 1970	0.167	0.131	0.295
Spain	Jan, 1970	0.123	0.104	0.224
Sweden	Jan, 1970	0.177	0.163	0.221
Switzerland	Jan, 1970	0.155	0.133	0.189
UK	Jan, 1970	0.158	0.129	0.246
USA	Jan, 1970	0.130	0.127	0.151
Average		0.146	0.123	0.229
Median		0.152	0.127	0.224

TABLE A2

SUMMARY STATISTICS: ANNUALIZED TOTAL RETURN IN U.S. DOLLARS

(EMDB DATABASE)

Country	Start	Arithmetic Mean	Geometric Mean	Standard Deviation
Argentina	Jan, 1976	0.571	0.204	0.926
Brazil	Jan, 1976	0.256	0.080	0.560
Chile	Jan, 1976	0.316	0.235	0.367
Colombia	Jan, 1985	0.326	0.200	0.295
Egypt	Jan, 1996	0.345	0.108	0.296
Greece	Jan, 1976	0.094	0.089	0.335
India	Jan, 1976	0.160	0.121	0.277
Indonesia	Jan, 1990	-0.057	-0.097	0.359
Jordan	Jan, 1979	0.100	0.081	0.168
Korea	Jan, 1976	0.116	0.113	0.325
Malaysia	Jan, 1985	0.065	0.037	0.288
Mexico	Jan, 1976	0.259	0.143	0.429
Nigeria	Jan, 1985	0.198	0.031	0.502
Pakistan	Jan, 1985	0.149	0.064	0.263
Peru	Jan, 1993	0.258	0.135	0.326
Philippines	Jan, 1985	0.292	0.196	0.365
Portugal	Feb, 1986	0.303	0.207	0.398
South Africa	Jan, 1993	0.171	0.096	0.224
SriLanka	Jan, 1993	0.041	-0.023	0.286
Taiwan, China	Jan, 1985	0.289	0.166	0.476
Thailand	Jan, 1976	0.125	0.095	0.297
Turkey	Jan, 1987	0.444	0.186	0.684
Venezuela	Jan, 1985	0.269	0.083	0.465
Zimbabwe	Jan, 1976	0.125	0.053	0.350
Average		0.217	0.108	0.386
Median		0.227	0.102	0.343

TABLE A3

SUMMARY STATISTICS: INSTITUTIONAL VARIABLES (FROM LA PORTA ET AL. (1998A))

Variable	Mean	Standard Deviation	Min.	Max.
Judicial Efficiency	8.311272	1.873544	3.25	10
Rule of Law	7.892043	2.281635	2.08	10
Corruption	7.688647	2.046631	2.92	10
Risk of Expropriation	8.674786	1.412053	5.22	9.98
Risk of Contract Repudiation	8.231078	1.618514	4.36	9.98
Accounting Standards	64.81173	10.03328	36	83
Antidirector Rights	2.979592	1.216035	0	5

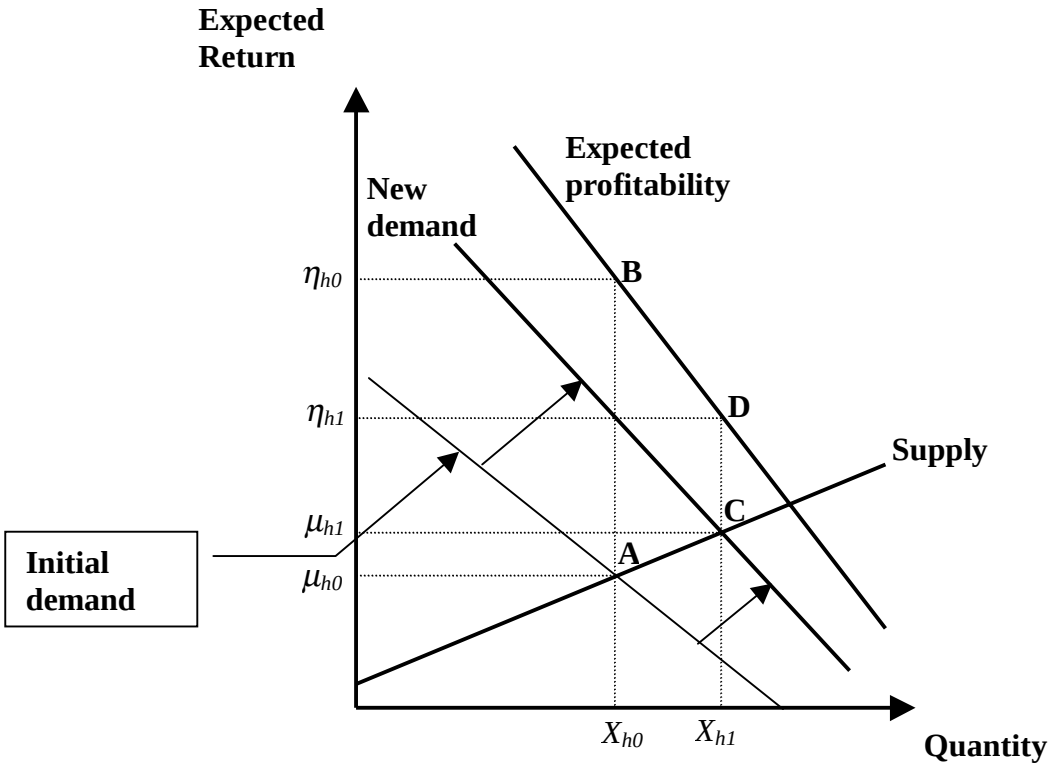

FIGURE 1

REDUCTION OF PRIVATE BENEFITS UNDER INTERNATIONAL SEGMENTATION

This figure depicts the effects of an improvement in the legal system that reduces the fraction of the company's profits that the manager can divert. After this improvement, managers can credibly commit to return more resources to outside investors: the demand for funds schedule shifts outward and the observed equilibrium point shifts from A to C. The rate of return increases from μ_{h0} to μ_{h1} , while the cost of equity capital decreases from η_{h0} (point B) to η_{h1} (point D). The equilibrium amount of equity finance increases from X_{h0} to X_{h1}.

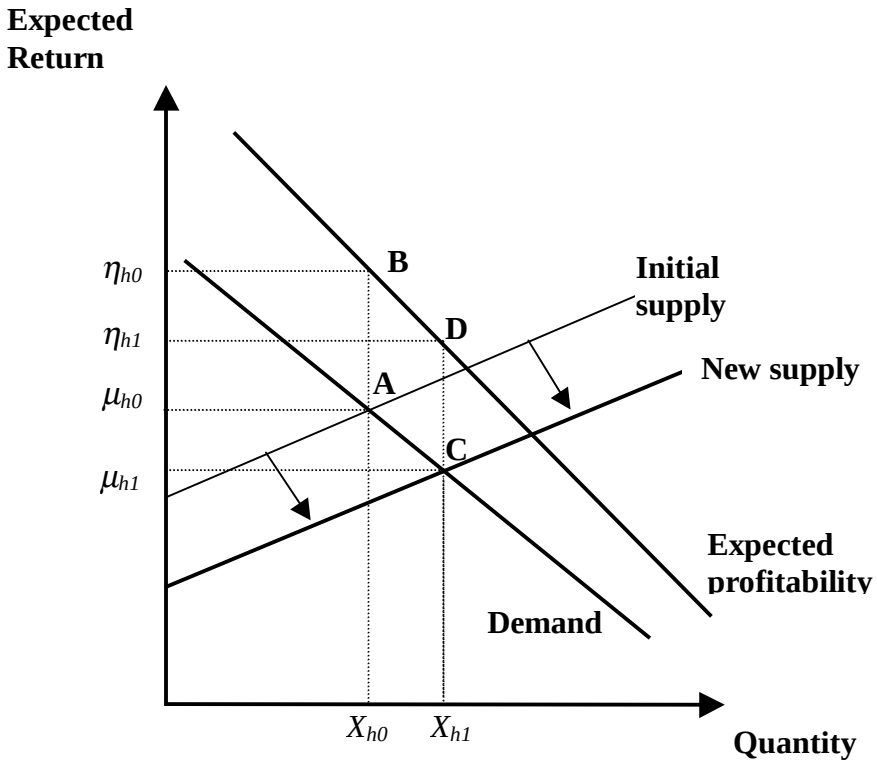

FIGURE 2

REDUCTION OF LEGAL AND AUDITING COSTS UNDER INTERNATIONAL SEGMENTATION

This figure shows the effects of a reduction in the legal and auditing costs that shareholder must bear to monitor managers. The investors' supply of funds schedule shifts downward and to the right, and the observed equilibrium point moves from A to C. The observed (expected) rate of return decreases from μ_{h0} to μ_{h1}, while the cost of equity capital to firms decreases, from η_{h0} (point B) to η_{h1} (point D). The equilibrium amount of equity finance increases from X_{h0} to X_{h1}.

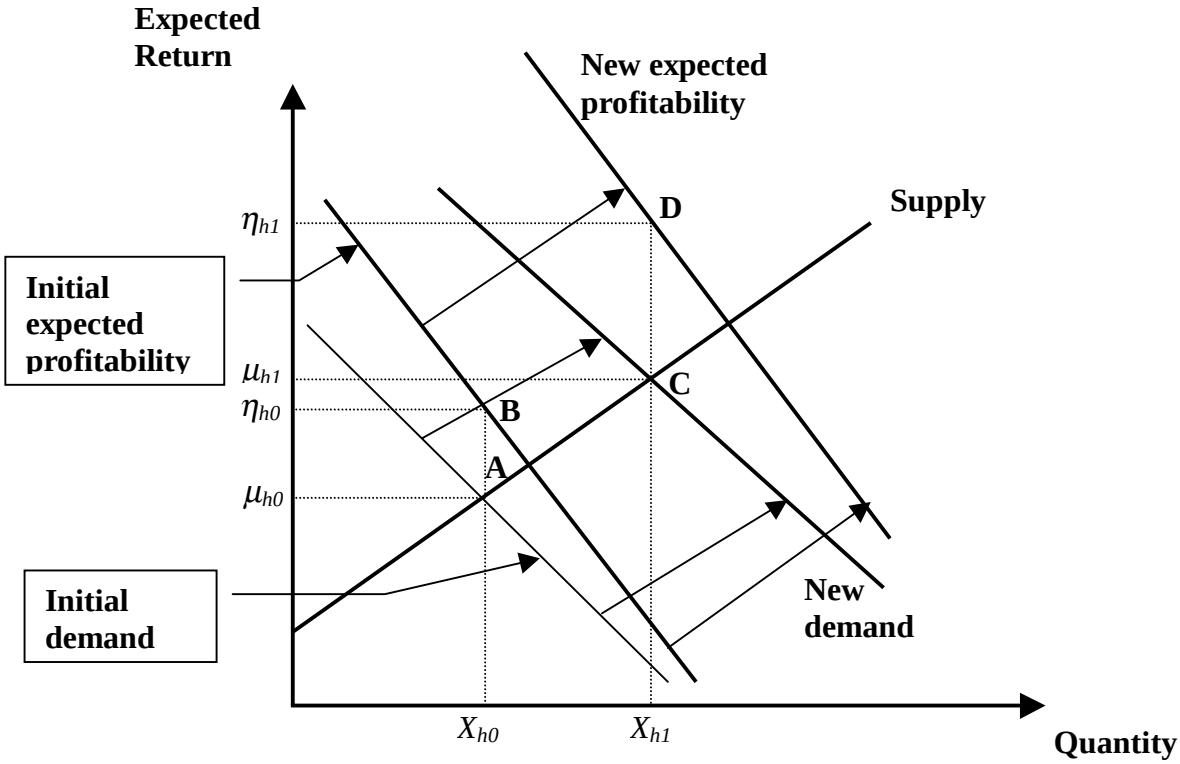

FIGURE 3

INCREASE IN PROFITABILITY UNDER INTERNATIONAL SEGMENTATION

This figure shows the effects of an improvement of the legal environment which increases the marginal productivity of capital. This is captured by an outward shift of the expected profitability schedule. The associated increase in the demand for equity capital shifts the observed equilibrium point from A to C. The observed (expected) rate of return increases (from μ_{h0} to μ_{h1}). The cost of capital increases from η_{h0} (point B) to η_{h1} (point D). The equilibrium amount of equity finance increases from X_{h0} to X_{h1}.

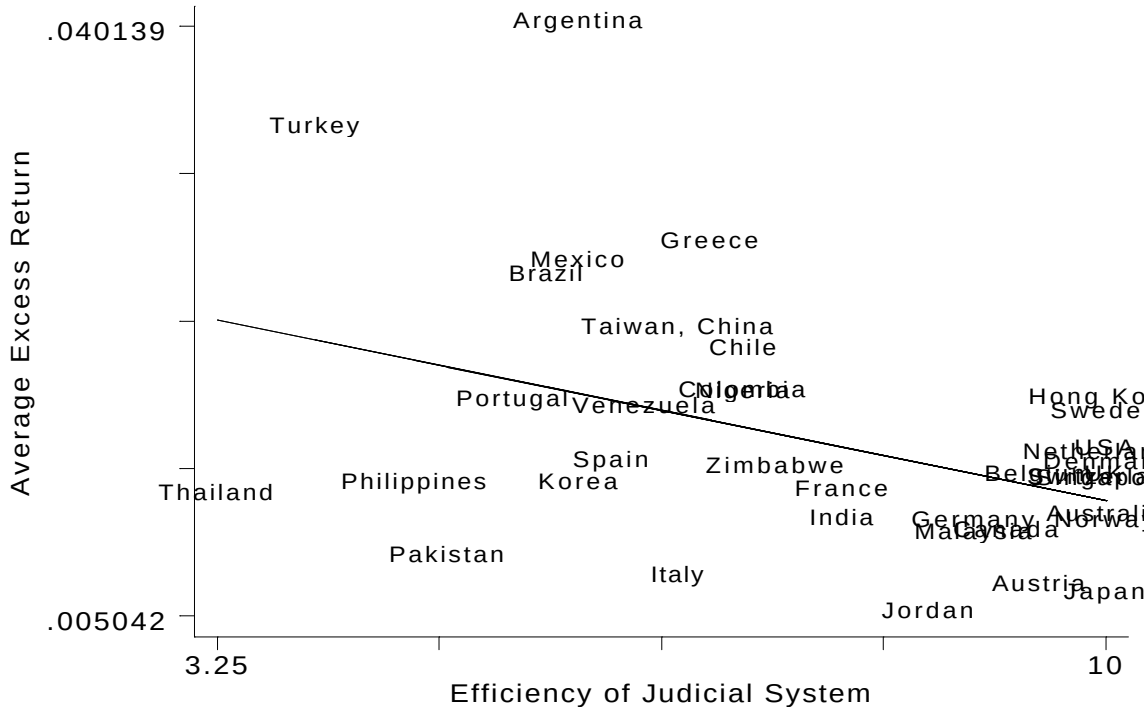

FIGURE 4

RETURNS VS. JUDICIAL EFFICIENCY– ALL COUNTRIES, 1987-1999

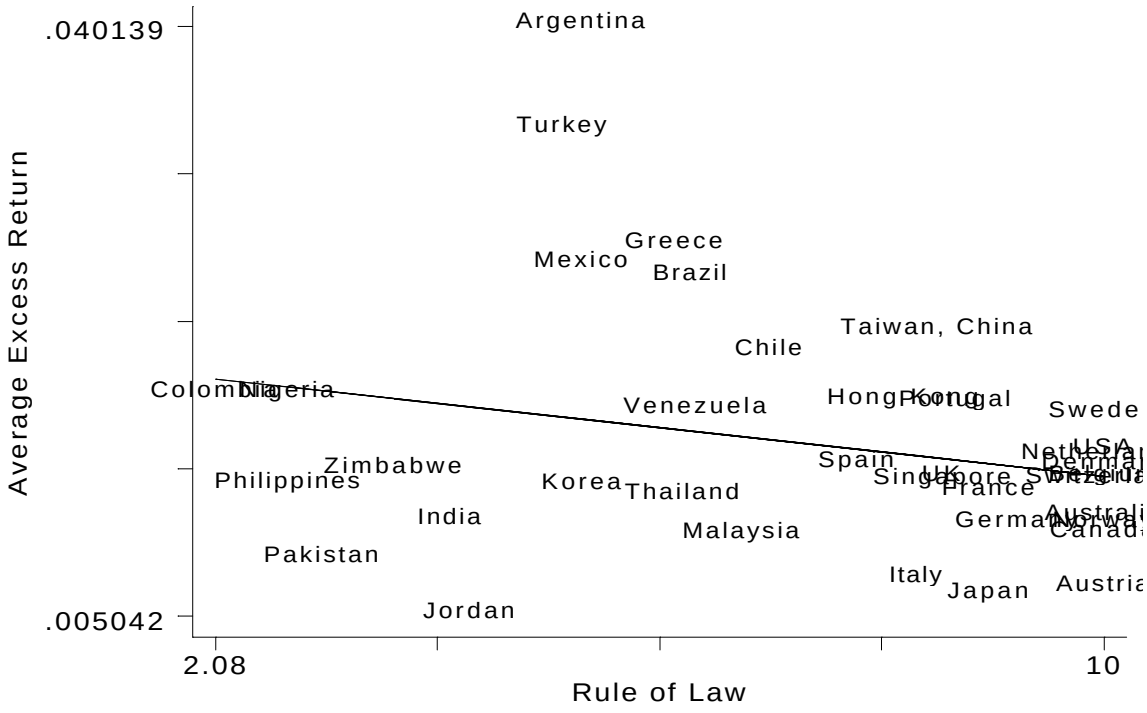

FIGURE 5

RETURNS VS. RULE OF LAW– ALL COUNTRIES, 1987-1999

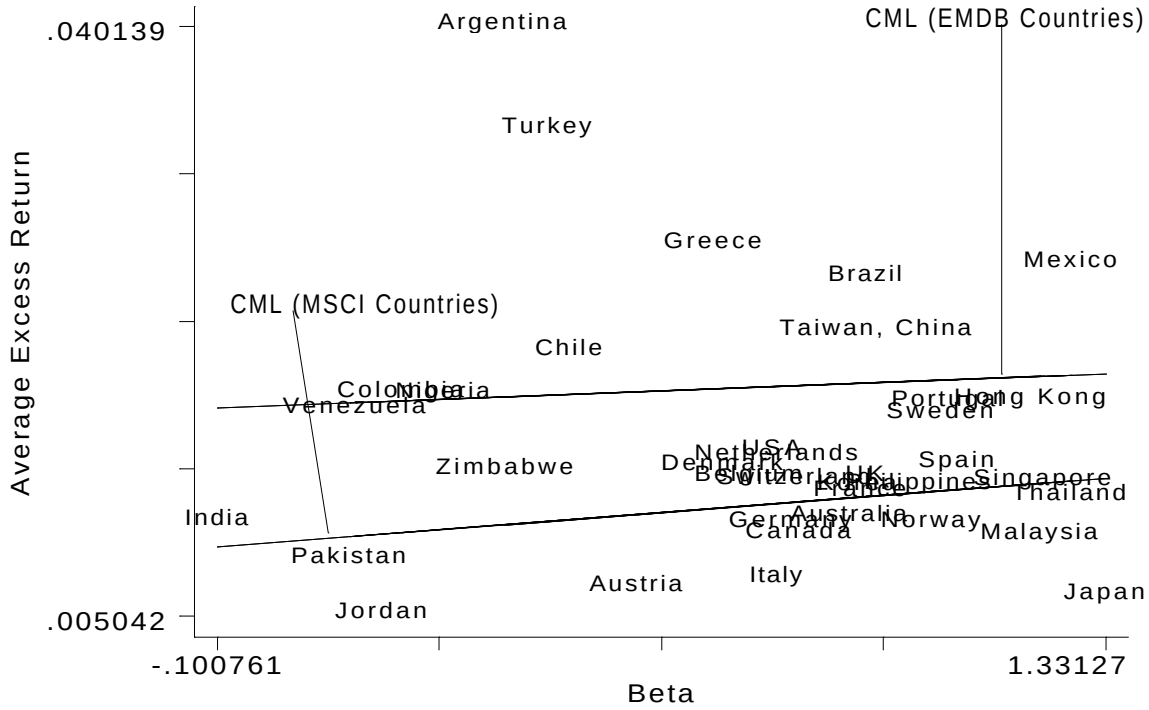

FIGURE 6

CAPITAL MARKET LINES (CML) – ALL COUNTRIES, 1987-1999

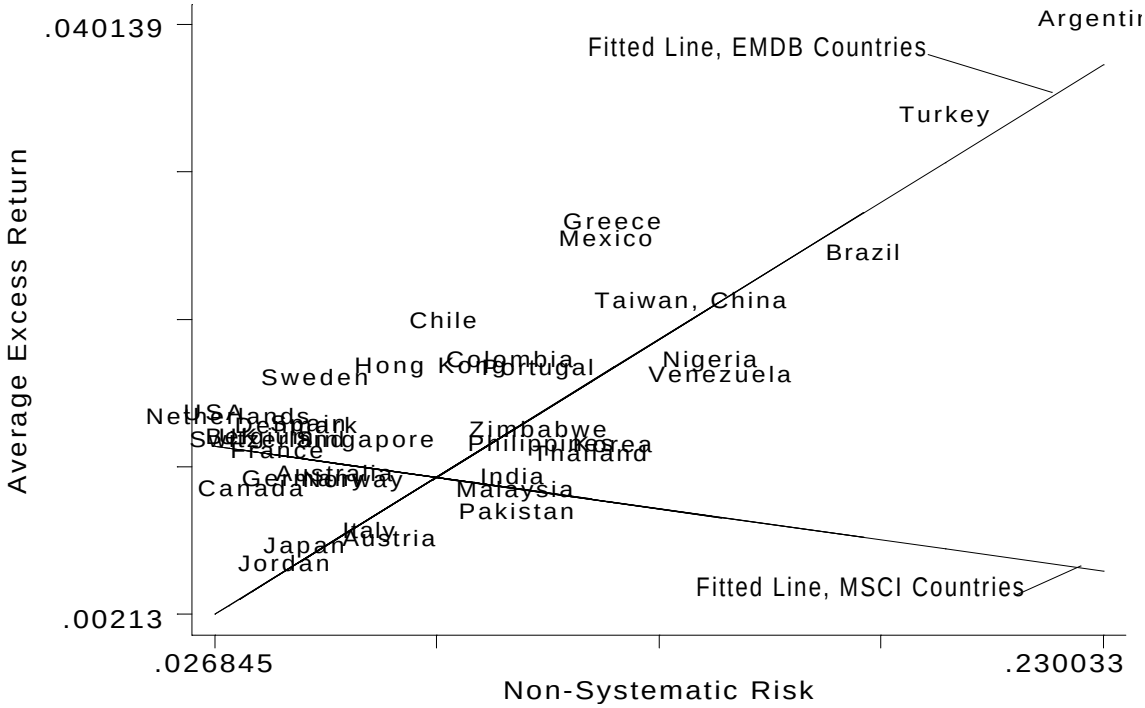

FIGURE 7

RETURNS VS. NON-SYSTEMATIC RISK – ALL COUNTRIES, 1987-1999

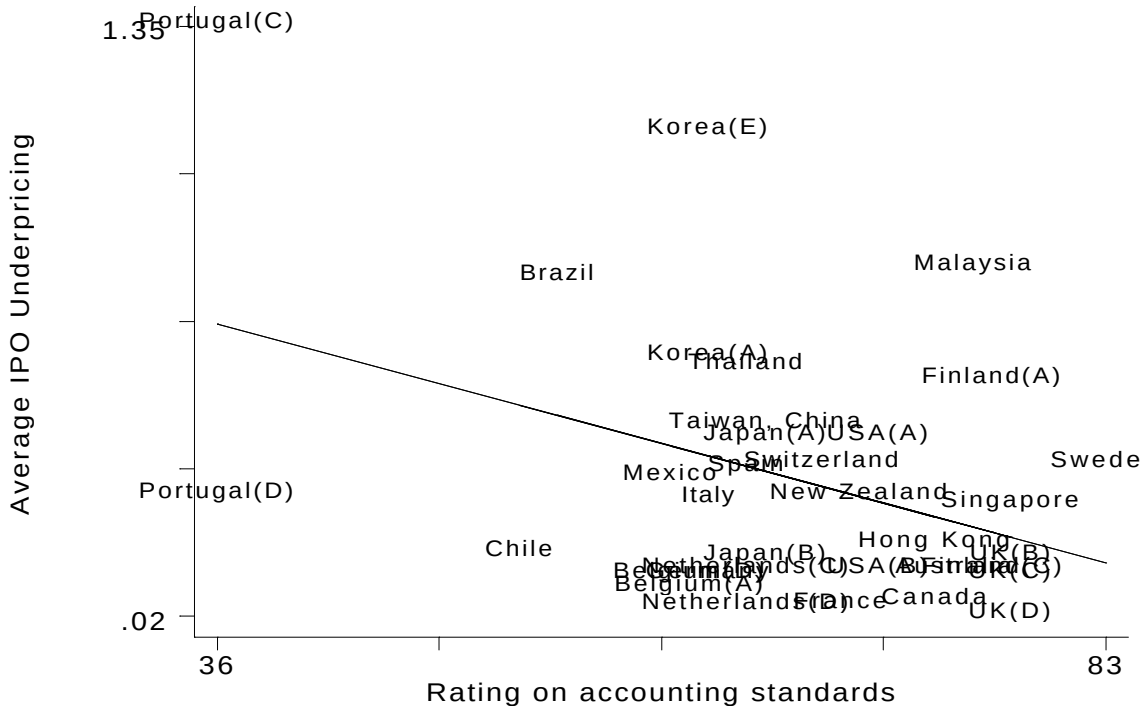

FIGURE 8

AVERAGE IPO UNDERPRICING VS. ACCOUNTING STANDARDS

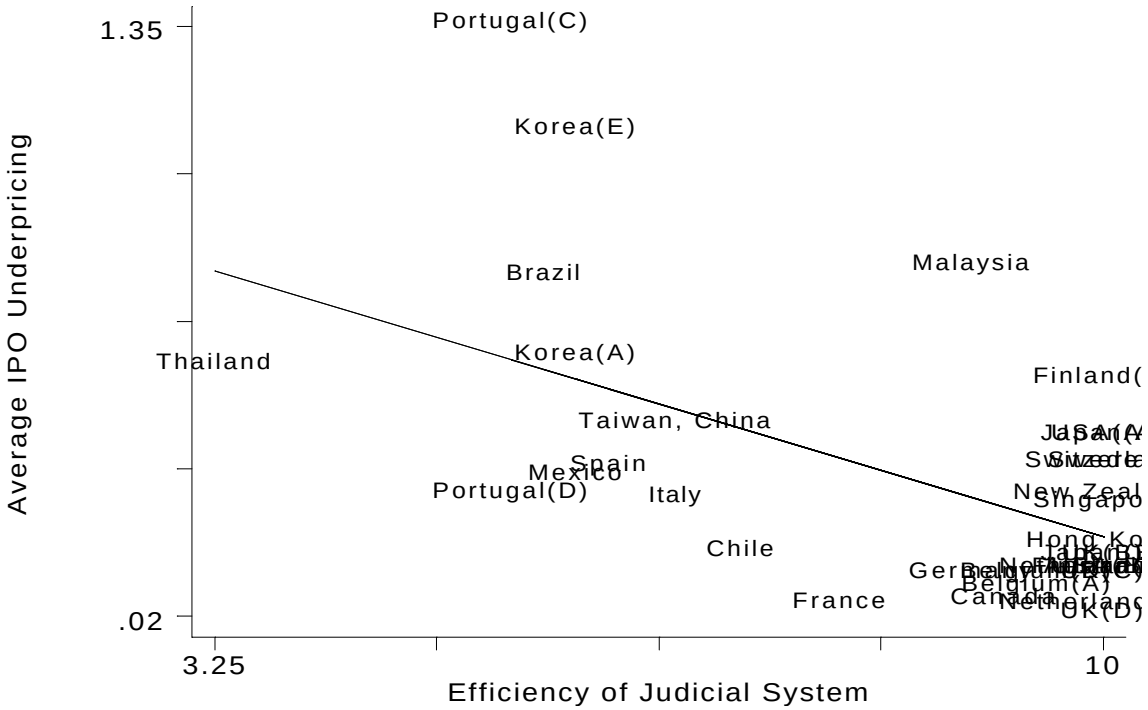

FIGURE 9

AVERAGE IPO UNDERPRICING VS. JUDICIAL EFFICIENC