Does Idiosyncratic Business Risk Matter?*

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Abstract

Several imperfections can prevent entrepreneurs from diversifying away the idiosyncratic risk of their business. As a result idiosyncratic risk discourages entrepreneurial activity and hinders growth, with the effects being stronger in economies with lower risk diversification opportunities. In accordance with this prediction, we find that OECD countries with low levels of risk diversification opportunities (as measured by the relevance of family firms or of widely held companies) perform relatively worse (in terms of productivity, investment, and business creation) in sectors characterized by high idiosyncratic risk. Differently from previous literature, we allow risk to be country specific. Since risk is endogenous to risk diversification opportunities, we instrument its value using sectoral risk in the US, a country where idiosyncratic business risk can be more easily diversified away. Tackling the endogeneity of risk and recognizing that it varies by country magnifies the estimated effects of risk on growth.

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1 Introduction

In standard Arrow-Debreu economies with complete markets, idiosyncratic risk can be fully diversified away and is irrelevant for equilibrium outcomes. But as emphasized by Townsend (1978) and Holmstrom (1979), among others, full risk diversification is costly and a great deal of theoretical research has analyzed how various financial frictions can prevent it, hampering aggregate productivity, output, and capital accumulation (Greenwood & Jovanovic, 1990; Bencivenga & Smith, 1991; Acemoglu & Zilibotti, 1997; and Meh & Quadrini, 2006). More recently Angeletos (2007) and Castro, Clementi & MacDonald (2004) have instead shown that the presence of undiversified risk can stimulate savings—because of either precautionary motives, or through an increase in entrepreneurial earnings—and can foster growth.

Despite much theoretical interest, there has been little empirical analysis of the effects of idiosyncratic risk on aggregate growth. A key issue in identifying the effects of idiosyncratic risk is that *observed risk*, typically measured by the volatility of some performance indicator, could be a very imperfect measure of the true *underlying risk* which determines agents' decisions. For example, principal agent models as in Holmstrom & Milgrom (1987) show that a trade-off generally exists between the benefits of risk-sharing and the provision of incentives. As a result the variability of agents' income (observed risk) only partially reflects the variability of exogenous shocks to output (underlying risk); see Guiso, Pistaferri & Schivardi (2005) for empirical evidence. In Thesmar & Thoenig (2010), an increase in firms' stock market participation endogenously decreases the observed risk of private firms and increases that of publicly traded firms. Thesmar & Thoenig (2000), Aghion, Angeletos, Banerjee & Manova (2010), and Bonfiglioli (2010) have also stressed that observed risk is endogenous to the market structure and to the risk diversification opportunities available in the economy.[1]

In this paper we provide evidence on the effects of idiosyncratic business risk on growth, accounting for the fact that observed risk is a distorted measure of underlying risk. To analyze this issue we consider a simple extension of the moral hazard model by Holmstrom & Tirole (1997) where risk averse entrepreneurs can choose projects with different risk-return tradeoffs. To solve the ensuing agency problem (which is a source of financial frictions), entrepreneurs have to partially finance business ventures with their own wealth and so id-

[1]Coles, Daniel & Naveen (2006), Fischer (2008), Bartram, Brown & Stulz (2009), and Panousi & Papanikolaou (2010) provide direct evidence that observed idiosyncratic risk is endogenous and is influenced by firms' decisions.

iosyncratic business risk cannot be fully diversified away.[2] Because of this, entrepreneurs may choose projects that are strictly dominated from the point of view of a well-diversified portfolio because they have a lower idiosyncratic risk. This hinders innovation, entrepreneurial activity and growth. The model delivers two key predictions that are common to a vast class of models: first, the effects of idiosyncratic business risk on growth strictly depend on existing risk diversification opportunities—with zero effects for big enough diversification opportunities and negative and increasing effects as diversifying risk becomes more difficult; second, the observed volatility of projects' returns is endogenous with respect to diversification opportunities, as entrepreneurs can endogenously reduce risk by choosing safer, more conservative projects.

To test the effects of idiosyncratic business risk on growth we use cross country-sector data for the group of OECD countries. Building on the methodology first introduced by Rajan & Zingales (1998), we consider a regression of sectoral growth on an interaction of the degree of country-level diversification opportunities and of sector-level idiosyncratic risk—after controlling for a full set of country and sector dummies and other time varying attributes. Theory predicts that countries with high diversification opportunities perform relatively better in sectors characterized by high idiosyncratic risk—i.e., the interaction term should be positive. To measure idiosyncratic risk, we follow Bulan (2005), Cummins, Hassett & Oliner (2006) and Panousi & Papanikolaou (2010) in thinking of the volatility of individual firms' stock market returns as a plausible measure for the risk of firms' activities. To decompose the overall observed variability of returns into a market and a firm idiosyncratic component we follow Campbell, Lettau, Malkiel & Xu (2001). As stressed above, a key problem in estimating the effects of risk on growth is that observed risk is endogenous to the diversification opportunities available in the economy. To tackle this issue, we instrument a country's sector-level idiosyncratic volatility with the analogous measure calculated in the US. The identifying assumption is that sectoral risk in the US is related to the risk that would have emerged in the other OECD countries if risk diversification opportunities were similar to those in the US. We allow the relationship to vary by country, since differences in fundamentals (for example in geographic and climatic conditions as well as in exogenous trade and technological patterns) can make idiosyncratic risk country specific. Castro, Clementi & Lee (2010) document that fundamentals affect risk.

In our model, differences in risk diversification are due to financial frictions but in

[2] Bitler, Moskowitz & Vissing-Jørgensen (2005) empirically document the contribution of entrepreneur's own wealth in solving agency problems. Hall & Woodward (2008) show that the idiosyncratic risk faced by entrepreneurs when starting new businesses is substantial.

practice several other factors, including cultural aspects or historical events, can limit entrepreneurs' ability to diversify away the risk of their businesses. To characterize the degree of diversification of business risk in the country and to identify its causal effects, we use direct indicators of the structure of business ownership as measured by the importance of family controlled firms or of widely held firms in the economy, as reported by La-Porta, Lopez-De-Silanes & Shleifer (1999) and Faccio & Lang (2002). In a country where firms are more family owned or less widely held, equity owners are less diversified and the economy is more sensitive to idiosyncratic business risk, which squares well with the evidence in Moskowitz & Vissing-Jørgensen (2002).

Our findings indicate that countries with low levels of diversification opportunities perform relatively worse in sectors characterized by high idiosyncratic risk. The implied effects are apparent only after explicitly accounting for the fact that observed and underlying risk differ. We also find that accounting for cross-country differences in idiosyncratic risk leads to larger estimated effects. Results are robust to the potential endogeneity of risk diversification measures and hold true with alternative measures of growth performance (in terms of labor productivity, capital, value added, and business creation) as well as of diversification. For example using, as in Dyck & Zingales (2004), more direct measures of private benefits of controls or other legal determinants of financial frictions, that can hinder risk diversification.

Our empirical approach extends the methodology of Rajan & Zingales (1998), which has been extensively used in the growth literature (see for example Pagano & Schivardi, 2003; Klapper, Laeven & Rajan, 2006; and Ciccone & Papaioannou, 2008). Typically the methodology hinges on assuming that the relevant sectoral characteristics (underlying risk in our case) are common across countries and that they can be directly computed from US data. Ciccone & Papaioannou (2007) extend the methodology to remove measurement error. Here we allow sectoral characteristics to vary across countries and show how to test for country specific differences. We find that taking them into account magnifies the effects of risk on growth.

Much literature has analyzed the relationship between idiosyncratic business risk and firm level performance. Among others, Bulan (2005), Cummins et al. (2006), and Caggese (2006) analyze the effects of risk on firms' investment and innovation activities, but without addressing the issue of the endogeneity of risk and firms' ownership structure. As in this paper, Panousi & Papanikolaou (2010) find that the effects of risk on firms' investment are magnified after addressing the issue. Castro, Clementi & MacDonald (2009) use the Rajan and Zingales' methodology with country-sector data to analyze the effect of idiosyncratic risk on sectoral employment size. Cũnat & Melitz (2010) and Manova (2007) use a similar approach but focus on the effects of risk on trade rather than on sectoral size. These three papers address the issue of the endogeneity of risk by assuming that underlying idiosyncratic risk is identical across countries and equal to the observed risk in the US. We show instead that there are country specific differences in idiosyncratic risk and we find that the effects of risk on growth are magnified by taking them into account.

The paper is also related to Koren & Tenreyro (2007), Jermann & Quadrini (2007), and Koren & Tenreyro (2008) who analyze how financial and technological progress affects both aggregate and idiosyncratic risk. In particular Koren & Tenreyro (2008) argue that, due to non convexities in the innovation process, more advanced economies can diversify business risk across a greater variety of products, which can explain why aggregate and idiosyncratic volatility both decrease with economic development. We have a different focus. First, our empirical analysis is based on OECD countries, mostly comprised of developed economies, for which differences in technology and in product variety might be limited. Second, we are interested in the causal effects of idiosyncratic risk on growth, which we identify using exogenous variation in risk across countries and sectors.

The remainder of the paper is structured as follows. Section 2 presents the theoretical framework. Section 3 introduces the empirical methodology. Section 4 describes the data. Section 5 presents results and discusses robustness. Section 6 concludes.

2 The model

We use a simple model to discuss how limited opportunities to diversify away idiosyncratic business risk affect productivity growth. The model emphasizes the distinction between the observed variability of firm returns—in brief *observed risk*—and the variability that would be observed under perfect risk diversification opportunities—what we call *underlying risk*. As in Holmstrom & Tirole (1997) agency problems induce financial frictions, which limit entrepreneurs' ability to diversify away the idiosyncratic risk of their business. Productivity growth is endogenous due to technological spill-overs as in the Schumpeterian paradigm reviewed in Aghion & Howitt (1998). The model yields two results, common to a vast class of models, which are important to identify the effects of idiosyncratic risk on growth. The first is that observed idiosyncratic risk is endogenous to the risk diversification opportunities of entrepreneurs. The second is that the effect of underlying idiosyncratic risk on productivity growth varies depending on risk diversification opportunities. The Technical Appendix contains the full details of the model. Here we present the model's main features and results more intuitively.

We consider a representative sector in an economy with n independent sectors that differ only in the level of underlying idiosyncratic business risk. The sector at time t is characterized by a given technological level, A_t, which determines the size of entrepreneurial projects. There is a measure one of entrepreneurs who live one period and after death are replaced by a new independent identical cohort. Entrepreneurs differ in terms of risk propensity, τ, which is a draw from the CDF F. Greater τ implies less curvature in the entrepreneur's utility function. Entrepreneurs can invest a fixed amount of capital A_t in either a risky or a safe project, with expected returns per unit of capital invested equal to μ_r and $\mu_s < \mu_r$, respectively. The safe project yields the return μ_s with certainty. The risky project yields a return λ with probability q and zero otherwise, so that its expected return is equal to $\mu_r = q\lambda$ and the *variance* of the project's return is given by

$$\sigma = \mu_r\lambda - \mu_r^2.$$

The parameter σ measures the *underlying idiosyncratic risk* in the sector: given μ_r, changes in σ have no consequences on the return of a well-diversified portfolio, but they can influence the choices of an undiversified entrepreneur. Increasing σ while keeping μ_r fixed means that the probability of a successful innovation is lower but its value in case of success is higher. This implies a greater idiosyncratic risk in innovation activities.

Entrepreneurial activity determines innovation and productivity growth as in Grossman & Helpman (1991) and Aghion & Howitt (1998). We assume that the growth rate of the technology level A_t is proportional to the number of successful risky projects with a factor of proportionality η per unit size of the innovation, as measured by λ.[3] Given that the constant over time population level is normalized to one, output and productivity are equal, and in equilibrium they both grow at the same rate as technology.[4] The resulting productivity growth rate is then equal to

$$\gamma \equiv \frac{\Delta A}{A} = \eta q \lambda(1-\rho) = \eta\mu_r(1-\rho) \tag{1}$$

where $(1-\rho)$ denotes the measure of entrepreneurs investing in the risky project.

[3]Here we are assuming that only successful risky projects induce a positive technological intertemporal externality but in practice we only need to assume that risky projects generate stronger spillovers than safer, more conservative projects.

[4]The model is such that the economy is always in a steady state; see Acemoglu & Zilibotti (1997), Acemoglu & Zilibotti (1999), and Koren & Tenreyro (2008) for models where the degree of financial market imperfections and market incompleteness depends on the development level of the economy.

The entrepreneur has just enough wealth to finance the project out of her own pocket. Still, she would like to sell the entire equity of the project and re-invests proceeds in a perfectly diversified portfolio, which bears no risk. But, as in Holmstrom & Tirole (1997), to solve moral hazard problems the entrepreneur has to retain a fraction β of the project equity, where β measures the severity of the moral hazard problem. So a fraction β of the project's idiosyncratic risk is born by the entrepreneur. When the moral hazard problem is more severe, the entrepreneur's ability to diversify away the idiosyncratic risk of her business is diminished. In equilibrium the entrepreneur will invest in the risky project if and only if her propensity to take risk is higher than a critical endogenously determined threshold τ^*. The resulting measure of entrepreneurs investing in the risky project is then given by $\rho = 1 - F(\tau^*)$, which is decreasing in β since τ^* is increasing in β. This is because with worse risk diversification opportunities (higher β), investing in the high-risk-high-return project yields lower expected utility due to the greater risk.

Since safe projects are riskless, the average variance of projects returns is equal to the product of the underlying idiosyncratic risk of a risky project σ times the share of entrepreneurs that invest in them, $(1-\rho)$. This means that

$$\omega = (1-\rho)\,\sigma \tag{2}$$

measures the *observed* average idiosyncratic risk in the sector. In the Appendix we prove that, when financial frictions are low enough (β is close to zero), all entrepreneurs invest in the risky project ($\rho = 0$) and observed and underlying risk are equal ($\omega = \sigma$). But, when β is sufficiently high, observed risk is imperfectly related to underlying risk and the derivative $\partial\omega/\partial\sigma$ is strictly less than one. In other words we have that:

Proposition 1 *In economies with high risk diversification opportunities (low β) the observed level of idiosyncratic risk ω accurately measures the underlying idiosyncratic risk σ. When risk diversification opportunities are low, observed risk is endogenous and moves less than one-for-one with underlying risk. The attenuation effect is larger the lower the risk diversification opportunities.*

In the Appendix we also prove that an increase in underlying idiosyncratic risk has a negative impact on productivity growth, $\frac{\partial\gamma}{\partial\sigma} \leq 0$. This is because greater σ leads to a reduction in the share of entrepreneurs who invest in the high-risk-high-return project. Further, we prove that the effect is stronger the worse the risk diversification opportunities, $\frac{\partial^2\gamma}{\partial\beta\partial\sigma} < 0$. This is intuitive since idiosyncratic risk reduces productivity growth only if risk diversification opportunities are low enough. In brief we have that:

Proposition 2 *An increase in underlying idiosyncratic risk σ reduces productivity growth γ. The effect is stronger the worse the risk diversification opportunities (larger β).*

Consider now the n independent sectors, each characterized by a different level of idiosyncratic business risk, denoted by σ_j, $j = 1, ..., n$. Sectoral differences in business risk might be due to differences in technology, in the degree of competition, or in the riskiness of innovation activities. Entrepreneurs have project opportunities in one sector. Sectoral productivity growth depends on the number of successful risky projects within the sector, exactly as in (1). The key assumption is that an innovation induces stronger technological spillovers within the sector than across sectors. Now consider two countries that differ in risk diversification opportunities β. Proposition 2 predicts that the country with worse diversification opportunities will grow relatively less in sectors with greater idiosyncratic risk. We can therefore relate the growth performance of a sector within a country to the corresponding level of idiosyncratic risk in the sector and then check how the relationship differs for countries with different risk diversification opportunities. In terms of the model this amounts to checking how $\partial\gamma/\partial\sigma$ differs for countries with different β, which measures the sign and magnitude of the second order partial derivative $\frac{\partial^2\gamma}{\partial\beta\partial\sigma}$. Based on this logic, we test whether sectors with higher idiosyncratic risk perform relatively worse in countries with less risk diversification opportunities. The empirical challenge is that observed risk ω is endogenous (see Proposition 1). As discussed in more details in the Technical Appendix, failing to recognize this leads to an important endogeneity bias.

3 Empirical methodology

The empirical test is based on the following regression model:

$$\gamma_{ji} = a_0 + a_1\beta_i \cdot \sigma_{ji} + a_2' X_{ji} + u_{ji} \tag{3}$$

where γ_{ji} is productivity growth of sector j in country i, β_i is a measure of the lack of diversification opportunities in country i (a proxy for β in the model), σ_{ji} is the level of underlying idiosyncratic risk in sector j in country i—i.e., the level of idiosyncratic risk that would be observed if risk diversification opportunities were sufficiently high (say for β sufficiently close to zero). Finally, X_{ji} are additional controls, including the log of the initial productivity level as well as sector and country dummies. Notice that the country dummy also captures possible technological spillovers across sectors. As in Rajan & Zingales (1998), the regression (3) identifies the effects of risk σ_{ji} on growth γ_{ji} by using within-country sectoral variability: for each country, we analyze how the relative performance of a sector varies depending on the corresponding relative level of idiosyncratic risk and we then analyze how the relationship differs for countries with different risk diversification opportunities. A negative a_1 indicates that sectors with higher idiosyncratic risk perform relatively worse in countries where the ownership structure is less diversified.

The problem in estimating a_1 is that we do not observe underlying risk σ_{ji} but only observed risk ω_{ji}. To solve this problem we model the relationship between ω and σ by assuming that

$$\omega_{ji} = c_{0i} + (1 - c\beta_i)\sigma_{ji} + \epsilon_{ji} \tag{4}$$

where $c > 0$. Here ω_{ji} is a measure of the observed idiosyncratic risk in sector j and country i and c_{0i} is a (possibly) country-specific term.[5] The positive c coefficient implies that an increase in underlying risk translates into a less than one-for-one increase in observed risk. The attenuation effect is stronger the lower the risk diversification opportunities (i.e., the higher is β). When risk diversification opportunities are high enough, observed and underlying risk are instead related one-for-one. This is a key implication of Proposition 1. Finally, ϵ_{ji} captures measurement error, which we assume to be independent from both β_i and σ_{ji}.

To estimate the model we need an instrumental variable condition. We assume that risk diversification opportunities in the US are large enough to guarantee that observed idiosyncratic risk is a relatively accurate measure of underlying risk.[6] Moreover we assume that underlying risk in a given sector in the US is related to the underlying risk in the corresponding sector in another country:

$$\sigma_{ji} = b_{0i} + \sum_{k=1}^{K} b_{ki} \left(\sigma_{jU}\right)^k + v_{ji} \tag{5}$$

where σ_{ji} is underlying risk in sector j in country i, b_{0i} is a constant term (that in principle is country specific), σ_{jU} denotes the level of underlying idiosyncratic risk in sector j in the US. This specifications allows for a non-linear relationship between underlying risk in country i and in the US ($K > 1$). The error term v_{ji} is assumed to be orthogonal to β_i and σ_{jU} at any power k. This specification encompasses several possibilities. One is that

[5]In a series of unreported exercises, we have also experimented with replacing the country dummies with a set of country characteristics that might determine overall idiosyncratic risk, such as GDP per capita, measures of financial development and reliance on oil exports. Results were very similar to those obtained using country dummies.

[6]As we will see in Section 4, our risk measures are based on stock market returns and it is well known that US firms stocks are owned largely by well diversified investors such as pensions funds, hedge funds, other institutional investors or households with broad portfolios (see for example Guiso, Haliassos & Jappelli, 2001 and La-Porta et al., 1999).

risk in the US is a perfect measure of the underlying risk in any country: $b_{1i} = 1$, $b_{ki} = 0$ for $k > 1$. This is the assumption made by Rajan & Zingales (1998) to identify the effects of financial market imperfections on growth. In this case, one can directly include the US measure of underlying risk in regression (3) and estimate the effects of risk on the economic performance. Of course, this may be a strong assumption, that we would like to test empirically. Here we allow for the possibility that underlying risk differs across countries and it is variably related to the risk in the US. In fact, countries differ in fundamentals: geographic and climatic conditions and exogenous technological patterns could differ, some countries entertain more direct economic relations with the US than others (say Canada versus Greece) or have institutions more similar to those prevailing in the US (such as common versus civil law countries). All this can affect underlying risk in the country and the degree of correlation with underlying risk in the US.

To estimate the cross term coefficient a_1 in (3) we can then use a Two-Stage-Least Squares procedure: we can estimate the coefficients b_{ki}'s, then replace σ_{ji} in (3) with $\sum_{k=1}^{K} b_{ki} (\sigma_{jU})^k$ and finally estimate equation (5) by OLS.[7] To estimate the coefficients b_{ki}'s we can use (5) to replace σ_{ji} in (4). This yields a regression model in terms of observables that can be estimated by Non-Linear Least Squares. To identify the coefficient c we need some independent variation in the level of risk diversification opportunities for countries that share the same relationship to risk as the US. Our identifying restriction is that the group of Scandinavian countries (Denmark, Norway, Sweden and Finland) satisfy this property, $b_{ki} = b_{kSCAND}$ for $i =$ Denmark, Norway, Sweden and Finland. This seems a plausible assumption, since these countries bear important similarities in sectoral composition, in geographical, social, and climatic conditions, etc.[8]

As is standard in the literature, we will measure observed risk with indicators of the idiosyncratic variability of stock market returns (Campbell et al. 2001). A less obvious choice is the measurement of diversification opportunities. In our model differences in risk diversification are due only to financial frictions. In reality, several other factors, including cultural aspects or exogenous shocks, can limit entrepreneurs' ability to diversify away risk. Given that we are interested in the causal effects of diversification of business owners on performance, rather than those of financial frictions *per se*, we start considering comprehensive, direct measures of diversification based on the structure of corporate own-

[7]Notice that if we were to replace σ_{ji} in (3) with $b_{0i} + \sum_{k=1}^{K} (\sigma_{jU})^k$ the result would remain unchanged, because the term b_{0i} would just become part of the country fixed effect that we include in equation (3).

[8]Of course identification requires some variation in the level of risk diversification opportunities within the group of countries, a condition that is satisfied by our data, as we show below.

ership in the country. Heaton & Lucas (2000) and Calvet, Campbell & Sodini (2007) show that entrepreneurs who obtain substantial income from their private business bear more idiosyncratic risk. Moskowitz & Vissing-Jørgensen (2002) document that owners of family businesses tend to hold a substantial portion of their overall personal wealth in the family firm and they are undiversified. So in a country where family businesses are diffuse, entrepreneurs bear more idiosyncratic risk and the importance of family firms (or conversely of widely held firms) can be used as a measure of the degree of diversification of entrepreneurs in the economy. These indicators are a more direct measure of the degree of diversification of business risk than indicators based on financial frictions alone. Of course, financial frictions do influence the ownership structure. For example, La-Porta, Lopez-De-Silanes, Shleifer & Vishny (1998) document that companies are less diversified and family firms are more common in countries with greater financial frictions while Burkart, Panunzi & Shleifer (2003) show how agency problems can account for the finding. In the analysis below we also consider some indicators of financial frictions which, as in our model, could be an important source of cross country differences in risk diversification opportunities.

A possible additional concern is that risk diversification measures could be endogenous to country specific differences in sectoral performance. For example, as the variability of sectoral performance increases, the incentive to diversify risk across sectors also increases, possibly leading to a change in the country's business ownership structure. Diversification measures could also be affected by measurement error, that in principle could be at least partly related to the relative performance of sectors in the country, which would then become a source of bias in the estimates. Although this seems less of a concern compared to the endogeneity of observed risk, we will tackle it anyway by instrumenting the diversification measure using changes in the demographic structure of the country's population induced by World War II. We think of WWII as an exogenous shock to the possibility to transmit businesses from one generation to the next and thereby to the degree of diversification of a country's business structure. These instruments exhibit important cross-country variation and they are most likely exogenous to the current sectoral performance of countries.

4 Data

We discuss the measure of idiosyncratic risk, sectoral performance, risk diversification and the instruments used. The Appendix provides further details.

4.1 Measures of idiosyncratic risk

Data on observed idiosyncratic risk for the OECD countries are calculated using information on monthly stock market returns from Thomson Datastream, which provides information on a large set of listed firms in 42 developed and emerging markets including all OECD countries.[9] To construct a country specific measure of sectoral idiosyncratic business risk, we closely follow the methodology in Campbell et al. (2001) and we decompose the overall volatility of yearly returns into a market and a firm component. The market is country specific.[10] Observed idiosyncratic risk is measured by the firm volatility component, which is computed for 36 sectors defined using the sectoral classification of Fama & French (1997).[11]

For the US we have two different sources of data to calculate idiosyncratic risk. We can use Thomson Datastream, as for all other countries, or we can take directly the indicator of idiosyncratic risk constructed by Campbell et al. (2001), which is based on CRSP data. As stressed by Ince & Porter (2006), measurement error is smaller in the CRSP data than in the Thomson Datastream data. To increase the power of the instruments we use both risk measures. Indeed the correlation between the Campbell et al. (2001) measure of risk and the analogous measure from Datastream is around 80 percent, which suggests that there is some independent variation, that can be exploited to improve estimation efficiency.

Table 1 reports the average, across sectors and years, of the value of idiosyncratic volatility in each country. Values vary from around .005 to 0.015 and are in the range of values computed by Campbell et al. (2001) for the US (see the last row in the table). As in Castro et al. (2010) we also find substantial cross-sectoral variation, indicating that sectors do differ in terms of observed risk. Sectoral coverage varies across countries (see last column in Table 1), although in most countries we have data for at least 20 sectors. An exception is New Zealand, for which only three sectors are available.

[9]Selection in the stock market differs across countries, which could lead to a bias in the measures of idiosyncratic volatility. See, for example, Thesmar & Thoenig (2010) and Himmelberg, Hubbard & Love (2004) for empirical evidence. Because of the inclusion of country dummies in (4) and in (5), our methodology is robust to biases that equally affect all sectors in a country. Moreover, we instrument sectoral risk with US risk, which might further address other potential selection problems.

[10]This assumption can be criticized on the grounds that financial markets are integrated and investors can invest internationally. But in practice the choice of the relevant market matters little, because the variance of the idiosyncratic component is large and dominates the market return component.

[11]In principle we might want to allow for an industry component different from the firm component. We do not do this because, in the theory of Acemoglu & Zilibotti (1999), sector specific risk matters: with high sectoral variability financial contracts are less efficient because it is harder to separate firm-level risk from sector-level risk. In any case, we checked that results are little affected by this choice. This is because the firm component dominates the sectoral component.

4.2 Measures of sectoral performance

Our first specification will focus on labor productivity growth measured as value added per worker at the yearly level. Productivity data are from the OECD-STAN database. STAN covers all sectors of the economy (at two digits) since 1970, although coverage varies by country and it is more comprehensive in more recent years (see Table 2). The number of sectors covered is generally high (above 20), with some exceptions including New Zealand and Portugal for which we have data for just four and six sectors, respectively. We also consider several other alternative measures of sectoral performance. The growth rate of value added, investment and capital labor ratio still come from the OECD-STAN database, while the index of business creation is taken from Bartelsmann, Scarpetta & Schivardi (2005). See their paper for details. Table 2 reports descriptive statistics for average productivity growth for each country (excluding the US, which is not used in the regressions to avoid endogeneity problems induced by the volatility measure). Overall, average productivity growth is around 2% per year, with a minimum of .5% in New Zealand and a maximum of 3.2% in Finland.[12] In total, we have 428 observations on productivity growth at the country-sector level. Despite being from official sources, sectoral productivity might be subject to measurement error. For example, employment is calculated in terms of number engaged rather than in terms of full time equivalent. To eliminate the influence of outliers, in all regressions we exclude observations above and below the first and last percentile of the cross country-sector productivity growth distribution—which are equal to -3.1% and 23%, respectively. We will also experiment with different weighting schemes to reduce the potential effects of measurement error.

4.3 Diversification measures and their instruments

As argued above, our preferred indicator of lack of diversification is the importance of family firms in the economy. Data on business ownership are taken from La-Porta et al. (1999) and Faccio & Lang (2002) for Western Europe and Mexico; Gadhoum, Lang & Young (2005) for the US; Claessens, Djankov & Lang (2000) for East Asia.[13] Firms are defined as family controlled if a single family or an individual is the majority shareholder and owns

[12] These comparisons are just illustrative of the data and should not be taken as indicators of the country's overall performance, as average growth may refer to different periods and sectors in different countries. The country dummies in the regressions control for cross-country differences in average growth.

[13] The data refer to one point in time in the mid-nineties. Ideally, one would like to have a full time series for the ownership structure. In reality, this is not likely to be a major problem, as ownership structure is very persistent. For example, Giacomelli & Trento (2005) analyze the ownership structure of Italian firms in 1993 and 2003, finding very modest changes, most of which are due to the privatization process that occurred over that decade.

at least 20% of shares. As an alternative measure of risk diversification opportunities of businesses owners, we consider the share of widely held firms in the economy, defined as those where there is no shareholder who owns more than ten per cent of the shares. Due to data constraints, these papers focus just on listed companies. Our assumption is that their ownership structure reflects economy-wide properties.[14] We follow the methodology in Mueller & Philippon (2006) to improve the comparability of data across countries.

The instruments for risk diversification opportunities are based on World War II demographic changes in terms of military, civil, and Jewish casualties in relation to the size of the country's population before the start of the War in 1939. As previously discussed, the idea is that war-related casualties have affected the demographic structure of the population, and thereby the possibility of transmitting businesses from one generation to the next. We use data on Jewish casualties because they might characterize differential effects of the War on the country's demographic structure and on the intergenerational transmission of businesses. As argued in footnote 13, business ownership structure is very persistent, so the effects of the war are likely to be present 40 years later. We take data for total population in 1939 and war related casualties from Wikipedia.[15]

In our model, the strength of agency problems affects the possibility of diversifying away idiosyncratic risk. Given this, we also consider some legal determinants of private benefits of control (which are a source of agency problem), such as indicators of the quality of accounting standard, rule of law, and anti-director rights as calculated by Dyck & Zingales (2004); see their paper for details. This allows us to check if our results are robust to the specific measure of diversification opportunities used.

Table 3 reports descriptive statistics for the different risk diversification measures used. In general, family firms and widely held corporations are very common: they represent around 50 and 32 percent of the firms in the sample, respectively. There are also important cross country differences, that indicate for example that Mexico and Continental Europe are much less diversified than the UK, the US, and Japan. In general, the relevance of widely held corporations and family firms are closely correlated (the correlation coefficient is minus .88). The correlation matrix at the bottom of the table also shows that the family indicator is negatively related to all three indicators of private benefits of controls, consistent with the notion that agency problems are an obstacle to risk diversification.

[14]Of course, the share of family firms in the group of private and listed companies does differ. However, we only need to assume that a country with relatively more family firms among the listed corporations will have also relatively more family firms in the economy. This is reasonable because the ownership structure of private and listed companies is influenced by common country-specific factors.

[15]See `http://en.wikipedia.org/wiki/World_war_ii_casualties`.

5 Results

We start by discussing the results of the first stage of the estimation procedure, that yields a characterization of the relationship between idiosyncratic risk in the US and in other countries. We then turn to estimating the effects of risk on productivity growth in countries with different risk diversification opportunities. Finally, we discuss the results with the alternative measures of risk diversification and sectoral performance.

5.1 First stage

As discussed in Section 3, we instrument observed sectoral risk in a country with US sectoral risk. Our instrumentation procedure entails the joint estimation of the b parameters in equation (5)—that characterizes the relationship between underlying risk in the US and in other countries—and of the c parameter in equation (4)—that characterizes the relationship between observed and underlying risk within a country. As previously discussed, we use measures of idiosyncratic risk in the US from both Thomson Datastream and CRSP. To allow for possible nonlinear effects we model the relationship between US risk and country specific risk in equation (5) using a third degree polynomial ($K = 3$). To keep the problem computationally manageable we allow only the two linear terms coefficients, $k = 1$, to differ across countries. Further details of the estimation procedure are reported in Appendix C.

Table 4 presents the results. The estimated value for c is positive, indicating that underlying and observed risk do differ. The estimated value of .63 together with our risk diversification measure implies that for all countries underlying and observed risk are positively related. The b parameters in (5) are statistically significant for at least one of the two volatility indicators in 14 out of 20 cases. In the case of Portugal and New Zealand the estimated coefficients are doubtful either because of their somewhat implausible magnitude or because of their sign (i.e., they are both negative). This is probably due to the small number of available sectors, as well as of firms within each of them (see Table 1). We will check that our results are robust to the exclusion or inclusion of these countries. For the remaining countries, results square up with expectations. In particular, there is an almost exact one-for-one relationship between risk in the US and in the UK and Canada. A joint test that both coefficients of the volatility indicators are zero is rejected for 15 out of 20 countries; for Italy, the acceptance is marginal (p-value of 10%), while it is more clear-cut for Korea, Mexico, New Zealand and Austria. Overall there is evidence that sectoral idiosyncratic risk differs across countries, since a test strongly rejects the null that the b coefficients are equal across countries. Columns three and six of Table 4 report the correlation of the risk of different sectors in a country with US sectoral risks, for the two US measures.[16] The correlation tends to be high for most countries, especially with the Campbell et al. (2001) measure, but always well below one. The most noticeable exceptions are Korea and New Zeland, for which both correlations are negative, reflecting small or negative coefficients in the estimation procedure. Again, our results are robust to the exclusion of these countries. All in all, this evidence goes against the assumption commonly underlying the Rajan & Zingales (1998) methodology, that US sectoral measures can be applied as such to all other countries.

5.2 Main results

We now discuss the estimates of the a_1 coefficient in equation (3). We start by considering pure cross section regressions, with growth and risk computed as time averages for the available years. In all regressions, we also include volatility not interacted, the log of initial productivity and country and sector dummies. Since it is unclear whether standard errors should be clustered at the country or at the sector level, standard errors are computed using the Huber-White method, which is robust to heteroscedasticity of unspecified form.[17] Following the discussion above, we also start excluding New Zealand and Portugal. We show below that all these specification choices have no major effects on the estimates. Column 1 in Table 5 reports the results when using observed country-sector volatility from the Datastream data, interacted with the measure of the importance of family firms in the country. We find that the interaction term is marginally significant at 10% and positive. This result would go against the idea that idiosyncratic risk has stronger negative effects on growth in countries with lower risk diversification opportunities. As discussed in the Technical Appendix, a positive a_1 coefficient is likely to be due to the fact that observed risk is endogenous, and that countries with worse risk diversification opportunities also have lower underlying risk.

To tackle the problem of endogeneity of observed risk, one could follow Rajan & Zingales (1998) in assuming that observed risk in the US measures the level of underlying risk in any other country. Column 2 of Table 5 reports the results when using the US volatility measure by Campbell et al. (2001) based on CRSP data. Column 3 deals with the analogous

[16]Due to the inclusion of country dummies in (4) and (5), we cannot directly compare risk levels across countries. This variation is however irrelevant for estimation results, since country specific differences in risk common across sectors are controlled by the country dummies in (3).

[17]We have also tried clustering standard errors at the country or at the sector level. We generally find that standard errors become smaller, particularly so when clustering at the country level. In this respect the significance level of the reported estimates is somewhat conservative.

measure calculated from Thomson Datastream. With the Campbell et al.'s measure, the coefficient on the interaction between family ownership and volatility is negative (-7.98) and highly significant (standard error 1.95), indicating that sectors with a higher idiosyncratic risk experienced lower productivity growth in countries where the ownership structure is less diversified. To give a better appreciation of the economic effect, other things equal, these estimates imply that the average productivity growth differential between an industry at the 75^{th} percentile of the sectoral distribution of idiosyncratic risk (Textile) and one at the 25^{th} percentile (Insurance and pension funds) is .8% higher in Canada (which corresponds to the 25^{th} percentile of the family ownership distribution by country) than in Italy (which corresponds to the 75^{th} percentile). When we use the measure of US idiosyncratic risk calculated from Thomson Datastream, the coefficient is again negative but not statistically significant at conventional levels (see column 3). This could be the result of the substantial measurement error present in the Thompson Datastream volatility measure. To investigate this possibility we follow Ashenfelter & Krueger (1994) and instrument the volatility measure from Datastream with the analogous measure from CRSP. If the errors in the two measures are uncorrelated, this procedure eliminates any bias induced by measurement error. We obtain a point estimate substantially larger (again in absolute value) and significant at 5%, which confirms that measurement error in the Thompson Datastream data is significant.[18]

The previous estimation hinges on assuming that sectoral underlying risk is the same across countries. Column 4 in Table 5 presents the results when we use the country specific measure of underlying risk discussed in the previous subsection. The coefficient now increases significantly in absolute value relative to the value obtained when imposing that underlying risk is equal across countries (from 7.98 to 12.17). This would again be consistent with the idea that countries with less diversified ownership structures also have lower underlying risk. Repeating the same comparative static exercise as above, we obtain that the difference in productivity growth between a sector with the idiosyncratic risk at the 75^{th} percentile of the risk distribution (Machinery in Finland) and at the 25^{th} percentile (Transport and Storage in Australia) would be 1.3% higher in Canada than in Italy.

[18]As a further check, we have also constructed an additional instrument following the procedure proposed by Ciccone & Papaioannou (2007). The idea is to run a regression of the form $\omega_{ji} = d_i + d_j + \theta_j \times d_j \times \beta_i$ where d_i and d_j are country and sector dummies while θ_j is the industry-specific response to country level differences in diversification opportunities β_i. Then we used the estimated $\hat{d}_j$'s and $\hat{\theta}_j$'s to construct a (predicted) value of the volatility present in the country with the highest level of diversification opportunities in the sample (which is equal to .1). This is obtained by calculating $\hat{\omega}_j = \hat{d}_j + .1 \times \hat{\theta}_j$ (see Ciccone & Papaioannou (2007), pg. 454, for details). With this instrument, the estimates are similar to those in columns [2] and [3], but with higher standard errors.

5.3 Robustness checks

To account for the endogeneity of family ownership, we instrument its interaction with volatility using demographic changes induced by WWII, also interacted. The first-stage regression, reported in the lower part of Column 1 of Table 6, indicates a positive effect of the ratio of civilian casualties over the country's population in 1939 on family ownership and a negative effect of the ratio of Jewish and military casualties over the country's population, with all effects being highly significant. A possible interpretation is that civilian casualties capture the destructive effects of WWII on the production system, with more destruction leading to more business creation after the war. Higher casualties among military and Jews may instead indicate a more dramatic reduction in the relatively young portion of the population, which has made it more problematic to transmit family businesses from the War generation to the next one. The instruments pass the Sargan test (p-value .44), which does not signal any endogeneity problems; the Anderson canonical correlation Likelihood Ratio test indicates that the rank condition is also satisfied (p-value .000) while the Cragg-Donald F statistic does not suggest any weak instruments problem. The upper part of Column 1 presents the second stage results. The interaction coefficient now falls to minus 14.6. The increase in the negative effects of risk on growth is most likely due to the fact that part of the negative correlation between the indicator of the importance of family firms and country underling risk is endogenous, say due to the fact that with less underlying risk the demand for risk diversification falls. We have also experimented with alternative instruments. In particular Mueller & Philippon (2006) show that the quality of labor relations are an important determinant of family ownership. They also show that measures of the degree of confrontation between European liberal states and guilds and labor organizations in the 18th and 19th centuries, as constructed by Crouch (1993), are strong predictors of the importance of family ownership in the country today. When using these alternative instruments for the family ownership indicator we find minor differences in the estimated coefficients (a coefficient of minus 12.9 rather than 14.6).

Sectoral employment varies greatly across countries and sectors. It goes from a few hundreds workers in computers and office Machinery in Greece to more than 6 millions workers in construction in Japan. To downplay the contribution of small sectors, possibly more subject to measurement error,[19] in column 2 of Table 6 we present the results when weighting the observations with the country specific sectoral employment size. The coefficient drops

[19] A regression of the square of the growth rate of productivity on log employment gives a coefficient of -.0032 with a standard error of .00086), indicating that smaller sectors have a more variable productivity pattern, possibly due to measurement error problems.

(in absolute value) to -8.5 and remains significant at 5%, indicating that our results are not driven by the behavior of small sectors. However, one problem is that this weighting scheme, although correcting for measurement error, might give rise to endogeneity problems. In fact, the patterns of sectoral specialization can be affected by the level of idiosyncratic risk—as emphasized by Cũnat & Melitz (2010) and Castro et al. (2009)—or more generally by the level of economic development, see Imbs & Wacziarg (2003). As an alternative, in column 3 we weight observations using a measure of the size of the country (total employment) multiplied by the sectoral shares of employment in the US, which are exogenous to the country's specialization patterns and can still correct for measurement error. We detect no substantial differences in the estimates with either weighting scheme. Only the fit of the regression improves substantially when weighting the observation with the US weights (the R^2 goes from .4 to .67). This indicates that, while not changing substantially the results, the regressions with exogenous weights fit the data better.

Another concern is that we only use a cross-section, taking the average of both productivity growth and volatility over a potentially long period of time.[20] Indeed, volatility might have changed substantially over time at the sectoral level, in which case an overall average might lead to misleading results. We address this issue in column 4 of Table 6, where we compute average productivity and volatility separately for each of the six five-year periods from 1973 to 2003. We then run the same regression as before by pooling all repeated cross sections. The results, reported in column 4, are very similar to those obtained with only cross sectional data, indicating that time aggregation does not bias our results.[21]

So far we have excluded New Zealand and Portugal from the sample, since their first stage results were somewhat dubious (see Table 4). Column 5 of Table 6 shows that the results are unchanged when including them in the sample.

One final concern relates to the possibility that our measures might be correlated with other characteristics of the financial system, beyond diversification and risk. In particular, Rajan & Zingales (1998) have shown that sectors with a higher external financial dependence (as measured by the share of capital expenditures not covered by firm cash flows in the corresponding sector in the US) grow relatively faster in countries with a more developed financial system. It could be that idiosyncratic risk is correlated with external dependence

[20] This is sometimes mitigated by the fact that the STAN data are often missing in initial years of the sample.

[21] We have also experimented with time series regression with growth and volatility measures at the yearly level. In the regression we allow for two year lags in the effects of volatility on growth, since entrepreneurial choices and productivity may be sluggish to respond to changes in risk. We obtain a lower coefficient (-3.6, significant at 1%), possibly because, at yearly frequencies, US volatility might be a more noisy measure of volatility in other countries.

and risk diversification with financial development. In this case our estimated coefficients could simply reflect the effect isolated by Rajan and Zingales. To address this concern, we add to our regression the interaction between the measure of external dependence and financial development proposed by Rajan & Zingales (1998).[22] The correlation coefficient between external dependence and our measure of idiosyncratic risk is .24. The correlations between family ownership and the two measures of financial development used by Rajan & Zingales (1998)—the stock market capitalization over GDP ratio and the domestic credit to private sector over GDP ratio— are -.42 and -.23, respectively. In the last two columns of Table 6 we report the results of the basic regression when also including the interaction between external dependence and either market capitalization (column 6) or private credit (column 7). As external dependence is only available for manufacturing sectors, the number of observations drops substantially. Yet in either specification the coefficient on the interaction between risk and diversification remain significant at 10% and large in absolute value (-17). The interaction between external dependence and the financial development indicators are positive, as expected, but statistically insignificant. This might be due to two factors. First, we are considering only OECD countries, for which differences in financial development are substantially smaller than in the original Rajan & Zingales (1998) sample, which also includes developing countries. Second, they consider value added growth while here we focus on productivity growth: it might be that credit availability is more important for overall growth, while, as the model shows, diversification opportunities have effects on productivity growth, even when financial resources are available.

5.4 Alternative risk diversification measures

One possible concern is that all our results so far are based on the importance of family firms. Family ownership might matter for performance for other reasons than risk diversification. For example, Caselli & Gennaioli (2005) argue that the ownership structure affects the selection of managerial talent. Table 7 presents the results with the alternative risk diversification measures discussed in Subsection 4.3: the share of widely held firms, indicators of the quality of the rule of law, anti-director rights and accounting standards.[23] The first is an alternative measure of the ownership structure of firms; the other three are indicators of legal determinants of financial frictions in the economy. Note that now a higher value of

[22]Appendix A explains how we adapted their sectoral classification (based on ISIC) to ours (based on STAN) and other data details.

[23]To simplify comparison of results, the only difference relative to the previous tables is that the interaction term in equation (3) is computed using these alterative measures.

the indicator is associated with greater risk diversification opportunities, so the interaction term is expected to be positive. This happens to be the case with all four indicators, with only accounting standards turning out not to be statistically significant at the 10% level. We take this as an indication that the estimated coefficients are actually capturing the effects of risk diversification opportunities on growth.

5.5 Other measures of sectoral performance

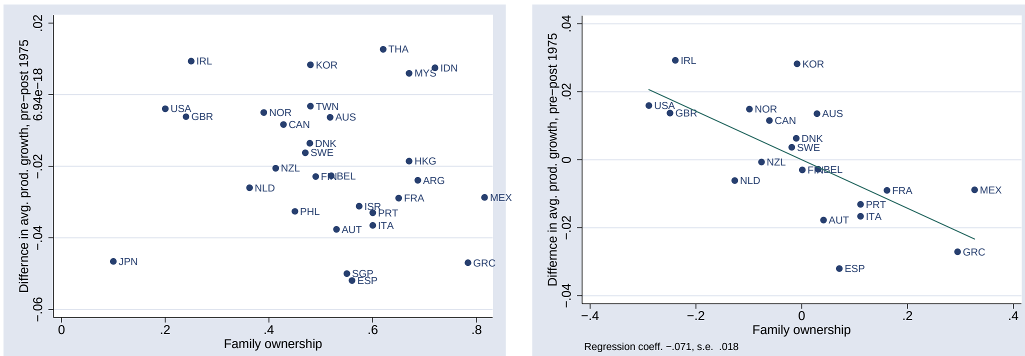
So far, we have focused the analysis on productivity growth. We now investigate whether our results hold true when we consider alternative measures of sectoral performance: value added growth, investment growth, capital-per-worker growth and business creation rates. The first is a natural measure of overall sectoral performance; the next two capture effects on capital accumulation; the last is an indicator of business creation. We expect that less opportunities to diversify risk will translate into lower investment and less business creation with greater effects in riskier sectors. The results appear in Table 8. Odd columns present the un-weighted regressions while even columns present results with US weights. Results with alternative weighting schemes are similar and are omitted to save space. We find that the effects of risk and diversification on growth tend to hold true also when considering these alternative measures of performance, although results are statistically significant only when weighting observations. This might be because the number of observations falls substantially when considering these alternative measures, reducing the precision of the estimates.

6 Conclusions

We analyzed the effects of idiosyncratic business risk on growth in OECD countries. We found that countries with low levels of diversification perform relatively worse in sectors characterized by high idiosyncratic volatility. The implied effects are apparent only after taking into account that observed and underlying risk differs and are magnified by taking into account that risk varies by country.

Our findings can shed some light on the diverging productivity performance of the US and Europe. Some authors (Ljungqvist & Sargent, 1998 and Comin & Mulani, 2006) have argued that the degree of idiosyncratic business risk has increased since the mid-1970s, due to an acceleration in the pace of technological progress and to the increased globalization of product markets. Our findings imply that countries with worse risk diversification opportunities should have experienced a more pronounced fall in productivity growth. Figure 1 displays some suggestive evidence that supports this hypothesis. In panel (a) we plot the difference in average productivity growth before and after 1975 against the share of family controlled firms. The figure characterizes a negative relationship between changes in

Figure 1: Productivity growth changes, pre-post 1975 vs. family ownership



(a) All countries (b) OECD countries, excluding Japan

Source: the Penn World Tables (Heston, Summers & Aten 2006)

productivity growth and the importance of family firms, with the exception of Japan, which went through a major depression, arguably unrelated to idiosyncratic risk and ownership structure. In panel (b) we restrict the sample to the group of OECD countries excluding Japan. The data line up nicely along a negatively sloped line. The large continental European economies, where family firms are more common, recorded a marked decrease in productivity growth. Productivity growth has instead accelerated in Anglo-Saxon countries, where the ownership structure allows better risk diversification. Our analysis provides a causal interpretation for this evidence. It suggests that the interaction of the increase in idiosyncratic risk with differences in risk diversification opportunities across countries can explain part of the US-Europe gap in productivity growth that has emerged over the last two decades, as well as the important differences across European countries. Investigating this issue further is an exciting avenue for further research.

A Sectoral concordance procedure

We construct the volatility measure for 49 different industries, following the industry classification of Fama & French (1997), which is also used by Campbell et al. (2001). STAN use the ISIC revision 3 sectoral classification, while Thomson Datastream use the ICB industry classification at the four digit level. Unfortunately, this does not match exactly with the industry classification used by Fama and French (FF). The table on the next page provides the sectoral concordance used to link the three classifications. In some cases, it was not possible to find a satisfactory correspondence for sectors; in some others, we were forced to aggregate sectors to achieve concordance across classifications. Specifically:

1. The following FF sectors had no clear correspondence in STAN or in Thomson Datastream and were dropped: toys (FF classification 6); motion pictures, amusement and recreation services (7); consumer goods (9); construction materials (17); fabricated products (20); precious metals (28); and shipping containers (40).
2. We aggregated the following FF sectors to match a corresponding sector in STAN and Thomson Datastream: food, soda and beer (FF 2, 3, 4); measuring equipment and medical equipment (12, 38); and defense, spacecraft, and aircraft (25, 27).
3. Four STAN sectors had no clear correspondence in FF and were dropped: fishing (STAN 05); wood and cork, excluding furnishing (20); other non-metallic mineral products (Thomson 266); and sales of motor vehicles (STAN 50).
4. The following Thomson Datastream sectors had no clear correspondence in FF and were dropped: recreational products (Thomson 3745); consumer electronics (3743); toys (3747); consumer goods (3767); gambling (5752); and recreational services (5755).

We ended up with a classification system based on 38 sectors, reported in the table below. In the regressions we also excluded personal services (34) and health care (11), as in many countries they are mostly provided out of the market (public provision, etc.).

To compute volatility, for each month, we take the observed return for each firm in the sample. For each country we then separately run a regression of firm returns on a full set of time dummies. The regression is weighted by using the previous period firm's market value. The residuals of this regression measures the firm's excess market return in the month. For each sector we then take the weighted average of the square of the residuals in a year where the weights are again the market value of the firm. This is our measure for the observed idiosyncratic risk of the sector in the given country and year, see Campbell et al. (2001) for further details.

Rajan & Zingales (1998) use the ISIC revision 2 classification system (restricted to manufacturing), while STAN is based on ISIC revision 3. We use a sectoral concordance table supplied by the OECD to match the two classifications. When one STAN sector corresponds to more than one ISIC sector, external dependence for the STAN sector is computed as a simple mean of its value in the corresponding ISIC sectors. The concordance procedure is reported in the "ISIC" column on the Sectoral concordance table below.

Sectoral concordance table

Fama French	STAN	Datastream	ISIC	Sector Name
1	01-02	3573		Agriculture
2,3,4	15	3533-7, 3577	311,313	Food and beverages
5	16	3785	314	Tobacco
8	22	5557	342	Printing and publishing
10	18,19	3765	322,323,324	Apparel and leather
11	85	4533		Health care
12,38	33	4537-73	385	Medical equipment
13	2423	4577	3522	Pharmaceutical
14	24ex2423	1353, 1357	3511,3513,352	Chemicals
15	25	3357	355,356	Rubber and plastic
16	17	3763	321,3211	Textile
18	45	1357, 1733, 2357, 3728		Construction materials
19	27	1753-7	371,372	Basic metals
21	29	573, 2753	382	Machinery
22	31	2733, 3722	383	Electrical machinery
23	36	2727,3724	332,390	Miscellaneous
24	34	3353-5, 2753	3843	Autos
25,27	353	2713,2717	384	Aircraft
26	351,352+359	2753	3841	Ships and railroad
29	13-14	1775		Mining of non energy prods.
30	10-12	1771		Mining of energy materials
31	23	533, 537, 577	353,354	Petroleum and natural gas
32	40-41	7535-77		Electricity, gas and water
33	64	5553, 6535-75		Post and telecom
34	80,90-93	5377		Personal services
35	71-74	2791-5, 2799, 5555, 9533-7		Other business activs.
36	30	9572-4	3825	Office equipment
37	32	2737, 9576-8	3832	Electronic equipment
39	21	1737	341,3411	Paper
41	60-63	2771-9, 5751, 5759		Transport and storage
42	51	2797, 5379		Wholesale trade
43	52	5333-75		Retail
44	55	5753, 5757		Hotel and restaurants
45	65	8355, 8773, 8779		Financial intermediation
46	66	8532-75		Insurance and pension funds
47	70	8733		Real estate
48	67	8737-71, 8775-7, 8985-95		Auxiliary to finance
7,9,17, 20, 28,40	No match	No match		See text
No match	5,20,26,28,50	2753, 3726, 3767, 5752		See text

B Ownership data and other diversification measures

La-Porta et al. (1999) compute their indicators considering only the largest 20 firms in each stock market, while the other papers discussed in the main text cover a much larger fraction of publicly traded companies. This latter approach is of course more informative, as the representation of large companies for the whole economy is limited. We will therefore use these indicators. For some countries, however, only the indicators based on the largest 20 firms are available. We follow Mueller & Philippon (2006) and we harmonize the data by running a regression of family ownership on comparable indicators of ownership structure using all countries where the data cover a large pool of companies. We then impute the value for the other countries by using the predicted values from this regression. Specifically, we regress the family ownership indicator based on the large fraction of firms on the fraction of medium-sized firms controlled by families, the fraction of value of top 20 firms controlled by families and the fraction of top 20 firms controlled by families, that are available for all countries. For countries for which the family indicator is missing (Australia, Canada, Denmark, Greece, Mexico,the Netherlands and New Zealand), we then use the predicted values from this regression. See Mueller & Philippon (2006) for further details.

C First stage estimation

To improve the relevance of instruments, we use data on the US idiosyncratic volatility from both Thompson Datastream and CRSP. Here we discuss how we generalize equation (5) to the case where both measures are used. To maximize degree of freedom, we also exploit time series variation. Equation (5) then becomes:

$$\sigma_{jit} = b_{0i} + \sum_{k=1}^{3}\sum_{z=1}^{2} b_{kiz}\left(\sigma_{jUzt}\right)^{k} + v_{jit} \tag{6}$$

where $z = 1$ indicates Thompson Datastream and $z = 2$ CRSP. Using (6) to substitute for σ in equation (4), we obtain:

$$\omega_{jit} = d_i + \sum_{k=1}^{3}\sum_{z=1}^{2} b_{kiz}(1 - c\beta_i)\left(\sigma_{jUzt}\right)^{k} + \eta_{jit} \tag{7}$$

where d_i captures any country specific effect and $\eta_{jit} = \epsilon_{ji} + (1 - c\beta_i)v_{jit}$, which is by assumption orthogonal to all independent variables. As explained in the text, to identify c we impose that the relationship between underlying risk in the four Scandinavian countries, (Denmark, Norway, Sweden and Finland) and in the US is the same. This leaves us with 17 different b's coefficients to be estimated for each regressor. To reduce the dimensionality of the estimation procedure, we impose that the quadratic and cubic terms are common across countries: $b_{kiz} = b_{kz}$ for $k > 1$. We then estimate equation (7), which involves a nonlinear estimation problem with 59 parameters to be estimated (i.e., c, 17 b_{1i1}'s , 17 b_{1i2}'s, b_{2i1}, b_{2i2}, b_{3i1}, b_{3i2}, the 19 country dummies b_{0i}'s and the constant). Note however that, conditional on c, the estimation becomes linear, as we can compute all terms $(1 - c\beta_i)\left(\sigma_{jUzt}\right)^{k}$. We therefore carry out the estimation using a line search method: we fix c, we compute the OLS estimates of the resulting linear estimation problem and we then search for the value of c that minimizes the residual sum of squares of the linear estimation. To implement the line search method we restrict the search for c over the range minus five to plus two and a half. This is reasonable since a value of c greater than two and a half would imply that more than three-quarters of the countries in the sample are on the negative side of the underlying-observed risk relationship, which may be regarded as highly implausible. The standard errors, in Table 4 are those of the linear estimation procedure; the standard error for c is instead calculated by bootstrapping. Finally, many countries miss observations for some sectors. To avoid loosing too many observations, we use out-of-sample fitted values, that is we calculate the measure of volatility also for sector-country observations for which no volatility is available from Thompson Datastream but it is available for the corresponding sector in the US data. For comparability with the cross-sectional growth regressions, the correlations in columns three and six of Table 4 use average risk over all available years. Results are similar when also using time series variability.

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Table 1: Volatility measures, descriptive statistics

Country	Mean	S.D.	N. sectors
AUS	.0055	.0020	26
AUT	.0064	.0029	16
BEL	.0076	.0050	20
CAN	.0142	.0249	30
DNK	.0053	.0018	11
ESP	.0089	.0050	28
FIN	.0136	.0061	20
FRA	.0088	.0079	32
GBR	.0122	.0220	31
GER	.0061	.0028	28
GRC	.0144	.0097	18
ITA	.0075	.0037	28
JPN	.0075	.0036	34
KOR	.0131	.0038	23
MEX	.0106	.0048	13
NLD	.0097	.0090	22
NOR	.0165	.0225	17
NZL	.0042	.0041	3
PRT	.0141	.0127	15
SWE	.0095	.0049	23
USA	.0066	.0026	35
USA (Camp.)	.0086	.0036	38

The table reports the cross-sectoral average volatility at the country level. Volatility of individual stocks is computed as the yearly standard deviation of monthly returns (net of the aggregate component). Sectoral volatility is the weighted average (according to market capitalization) of individual volatility. The last row reports the volatility computed by Campbell et al. (2001). See Subsection 4.1 and Appendix B for sources and definitions.

Table 2: Descriptive statistics for productivity growth, by country

Country	Mean	Median	S.D.	N. of sects.	First Year	Last Year
AUS	.015	.013	.021	7	1975	2001
AUT	.023	.016	.016	30	1977	2003
BEL	.020	.018	.018	9	1971	2003
CAN	.014	.009	.017	24	1971	2003
DNK	.025	.016	.020	33	1971	2003
ESP	.016	.009	.038	32	1981	2003
FIN	.032	.025	.021	33	1971	2003
FRA	.018	.021	.027	34	1979	2003
GBR	.016	.010	.020	10	1972	2003
GER	.016	.010	.029	30	1992	2003
GRC	.028	.026	.026	31	1996	2003
ITA	.016	.012	.018	26	1971	2003
JPN	.017	.020	.018	17	1971	2003
KOR	.028	.023	.026	5	1971	2003
MEX	.009	.017	.021	24	1981	2003
NLD	.010	.008	.015	23	1971	2003
NOR	.025	.031	.023	31	1971	2003
NZL	.005	.006	.030	4	1990	2002
PRT	.025	.017	.020	6	1978	2003
SWE	.024	.019	.025	19	1971	2003
Total	.019	.016	.022	428	1971	2003

The table reports descriptive statistics for average yearly productivity growth for the observations used in the regressions of Table 5 and 6. The data come from the OECD Stan database. Statistics are computed across sectors within country, using national sectoral employment as weight. "N. of sects." is the number of sectors for which data are available in a given country; "first" and "last year" are the first and last year for which productivity growth in any sector is available in a given country.

Table 3: Diversification measures

Country	Family Firms	Widely Held Firms	Rule of Law	Anti Direct. Rights	Accounting Standards
AUS	.52	.44	10	4	75
AUT	.53	.11	10	2	54
BEL	.52	.2	10	0	61
CAN	.43	.49	10	5	74
DNK	.48	.33	10	2	62
ESP	.56	.26	7.8	4	64
FIN	.49	.29	10	3	77
FRA	.65	.14	9	3	69
GBR	.24	.63	8.6	5	78
GER	.65	.1	9.2	1	62
GRC	.78	.12	6.2	2	55
ITA	.6	.13	8.3	1	62
JPN	.1	.8	9	4	65
KOR	.48	.43	5.3	2	62
MEX	.82	.074	5.3	1	60
NLD	.36	.23	10	2	64
NOR	.39	.37	10	4	74
NZL	.41	.34	10	4	70
PRT	.6	.22	8.7	3	36
SWE	.47	.39	10	3	83
USA	.2	.65	10	5	71

Correlation matrix

Family	1				
Wid. held	-.88	1			
Rule of Law	-.48	.23	1		
Ant. rights	-.62	.74	.30	1	
Acc. Stand.	-.43	.47	.36	.48	1

The table reports the values for the diversification measures for the countries used in the regressions, as well as the correlation matrix. See Subsection 4.3 and Appendix B for sources and definitions.

Table 4: First stage regression

	Campbell			Datastream			$H_0: b_1 = b_2 = 0$
	b_1	S.E.	Corr.	b_2	S.E.	Corr.	p-value
AUS	0.42	0.48	0.18	1.12^b	0.45	0.72	.009
CAN	0.92^c	0.52	0.89	0.73	0.59	0.36	.016
GRC	1.97^b	0.77	0.69	-1.09	1.47	-0.29	.038
ITA	0.53	0.5	0.22	0.62	0.48	0.11	.116
JPN	0.58	0.45	0.52	0.73^c	0.39	0.40	.020
KOR	0.07	0.55	-0.92	0.58	0.68	-0.28	.636
MEX	0.52	0.84	0.46	3.39	2.63	0.84	.366
NZL	-0.02	10.44	-0.70	-1.48	71.29	-0.47	.998
PRT	10.21^a	1.11	0.55	-14.33^a	1.97	-0.46	.000
AUT	0.82	0.85	0.37	-0.02	1.09	-0.53	.447
BEL	0.69	0.51	0.76	0.76	0.47	0.43	.034
ESP	1.06^c	0.58	0.85	1.38^c	0.84	0.61	.008
FRA	1.02^b	0.51	0.83	1.44^a	0.53	0.63	.000
GBR	0.98^b	0.46	0.90	0.84^b	0.4	0.43	.001
GER	0.65	0.5	0.70	0.86^c	0.47	0.55	.020
NLD	0.85^c	0.48	0.86	0.9^b	0.43	0.49	.003
SCAND.	1.00^b	0.49	0.84	1.29^a	0.44	0.60	.000
Square	-26.35	27.25		-45.68	26.17		.088
Cube	336.27	472.33		571.06	415.83		.280
		Coeff.		S.E.			
c^*		$.63^c$		.37			
R^2			0.06				
N. obs.			5,834				

The table reports the estimated coefficients of equations (4) and (5) in the main text. We allow for a country specific linear coefficient on the US measure of idiosyncratic volatility from Campbell et al. (2001) b_1 and from Thomson Datastream b_2, with the exception of the Scandinavian countries (Denmark, Finland, Norway and Sweden) for which a common coefficient is imposed. We also include common quadratic and cubic terms. The standard error for c^* is obtained by bootstrapping. The columns under "Corr." report the correlation between the resulting measure of sector level idiosyncratic risk and the two US measures from CRSP and Datastream. See Appendix C for details.

Table 5: Productivity growth, diversification and idiosyncratic risk

	Volatility measure			
	Observed	US-CRSP	US-Datastr.	Underlying
	[1]	[2]	[3]	[4]
Family∗ volatility	2.84^c (1.55)	-7.98^a (2.84)	-6.47 (5.71)	-12.17^a (5.12)
Initial prod.	$-.020^b$ (.009)	$-.018^a$ (.006)	$-.019^a$ (.006)	$-.020^a$ (.006)
R^2	0.56	0.4	0.41	0.42
N. obs.	265	426	391	387

The dependent variable is the average growth rate of labor productivity at the sectoral level, computed over all the available years. Volatility is measured as: column [1], observed volatility, computed using Datastream; column [2], US volatility, computed by Campbell et al. (2001) using CRSP; column [3], US volatility computed using Datastream; column [4], underlying volatility computed using the procedure detailed in Section 3. All regressions include a full set of industry and country dummies. Robust standard errors in parenthesis. ^{a} indicates significance at 1%, ^{b} at 5%, ^{c} at 10%.

Table 6: Robustness checks

	[1]	[2]	[3]	[4]	[5]	[6]	[7]
Family∗	-14.65^b	-8.55^a	-8.56^a	-10.94^a	-9.70^a	-17.11^c	-17.43^c
volatility	(7.07)	(3.29)	(3.21)	(3.86)	(3.86)	(9.39)	(9.49)
Ext.Dep∗						-.040	-.054
Find. Dev.						(.031)	(.044)
Initial prod.	-.020^a	-.013^b	-.009^b	-.019^a	-.020^a	-.027^a	-.027^a
	(.006)	(.005)	(.004)	(.006)	(.006)	(.008)	(.008)
R^2	0.41	0.55	0.67	0.34	0.42	.42	.43
N. obs.	387	387	318	1010	395	213	213
Weight	NO	OWN	US	NO	NO	NO	NO
Repetead CS	NO	NO	NO	YES	NO	NO	NO
First Stage IV Regression							
Military	-2.26^a						
	(.519)						
Civilian	10.12^a						
	(1.13)						
Jews	-16.08^a						
	(3.40)						
LR stat. (p)	0.00						
Sargan (p)	0.44						
Cragg-Donald stat. (p)	0.00						

The dependent variable is average growth rate of labor productivity at the sectoral level, computed over all the available years. Volatility is the underlying volatility computed using the procedure detailed in Section 3. The first column reports the results from the IV regression where the interaction variable is instrumented with WWII casualties. Military, Civilian and Jews are casualties in each group over the population in 1939, interacted with the volatility measure. All other regressions are OLS. In column [2] we weight observations according to sectoral employment while in column [3] according to total country employment multiplied by the US sectoral share of employment. In column [4] we take 5 year averages of the variables (rather than a single cross section). In column [5] we repeat the basic regression of column [4] of Table 5 keeping the observations for Portugal and New Zealand in the sample. In columns [6] and [7] we add the interaction between external dependence and financial development measured by stock market capitalization over GDP in column [6] and private credit over GDP in column [7] as in Rajan & Zingales (1998). All regressions include a full set of industry and country dummies. Column 1 in the second panel reports the results of the first stage regression where family is instrumented using WWII casualties. Robust standard errors in parenthesis. ^{a} indicates significance at 1%, ^{b} at 5%, ^{c} at 10%.

Table 7: Other measures of diversification

	Widely held firms	Rule of law	Anti directors rights	Accounting standards
	[1]	[2]	[3]	[4]
Diversif.∗ volatility	12.98^c	1.16^a	1.09^c	.21
	(7.3)	(.46)	(.62)	(.14)
Initial prod.	$-.020^a$	$-.020^a$	$-.020^a$	$-.020^a$
	(.006)	(.006)	(.006)	(.006)
R^2	0.41	0.42	0.41	0.41
N. obs.	387	387	387	387

The dependent variable is average growth rate of labor productivity at the sectoral level, computed over all the available years. Volatility is the measure of underlying risk computed using the procedure detailed in Section 3. Widely held firms is the share of listed firms that are widely held. Rule of law, anti-director rights and accounting standards are the determinants of the Private Benefits from Control (PBC) proposed by Dyck & Zingales (2004); higher values of the indicators imply higher investor protection and lower PBC. Robust standard errors in parenthesis. ^{a} indicates significance at 1%, ^{b} at 5%, ^{c} at 10%.

Table 8: Alternative performance measures: growth rate of value added, of investments, of investment per worker, entry rate

	Value added growth		Investment growth		Capital per worker growth		Entry rate	
	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]
Family∗ volatility	-1.59	-6.59[b]	-11.11	-32.18[a]	-19.98	-29.40[a]	-4.38	-9.92[c]
	(2.67)	(2.62)	(11.81)	(9.43)	(13.94)	(8.82)	(7.00)	(5.33)
Initial level	-.036[a]	-.021[a]	-.002	-.007[b]	-.018[c]	.013[b]	-.007[b]	0.17
	(.003)	(.003)	(.010)	(.003)	(.009)	(.005)	(.003)	(.15)
- R^2	0.64	0.75	0.28	0.78	0.28	0.65	0.84	0.77
N. obs.	436	360	265	219	262	217	131	121
Weight	NO	YES	NO	YES	NO	YES	NO	YES

The dependent variable is average growth rate of real value added in columns 1-2, of real investment in columns 3-4, of investment per worker in columns 5-6; in the last two columns is average entry rate. All variables are at the sectoral level, computed over all the available years. The weights are computed as total country employment multiplied by the US sectoral share of employment. Robust standard errors in parenthesis. [a] indicates significance at 1%, [b] at 5%, [c] at 10%.

A Technical Appendix: detailed derivations

The model informally discussed in Section 2 in the main text builds on Holmstrom & Tirole (1997). Growth is endogenous due to technological spillovers as in the Schumpeterian paradigm reviewed in Aghion & Howitt (1998). For the sake of exposition, the model is simplified so that the economy is always in a steady state; see Acemoglu & Zilibotti (1997), Acemoglu & Zilibotti (1999), and Koren & Tenreyro (2008) for models where the degree of financial market imperfections and market incompleteness depends on the development level of the economy.

A.1 Assumptions

To keep notation simple we consider a representative sector in an economy with n independent sectors that differ only in the level of underlying idiosyncratic business risk. The sector at time t is characterized by a given technological level A_t, which determines the size of entrepreneurial projects. There is a measure one of entrepreneurs who live one period. Entrepreneurs born at time t have an initial amount of wealth equal to A_t and quadratic consumption preferences:

$$E\left[U_t(c)\right] = E\left(c - \frac{1}{\tau A_t}c^2\right).$$

Entrepreneurs differ in terms of risk propensity τ, which is draw from a uniform distribution with support $[\underline{\tau}, \bar{\tau}]$. Moreover, the propensity to take risk is scaled by the state of technology A_t to guarantee that the coefficient of relative risk aversion is constant, a necessary condition for a balanced growth equilibrium. We would obtain identical results by postulating a Constant Relative Risk Aversion utility function and then take a second order expansion around the steady state equilibrium. Entrepreneurs live one period and after death they are replaced by a new independent identical cohort.

Entrepreneurs can invest in a project that costs A_t unit of wealth. Projects could be risky or safe, with expected returns per unit of capital invested μ_r and $\mu_s < \mu_r$, respectively. Project choice is irreversible. The safe project yields $\mu_s A_t$ with certainty while, if the entrepreneur behaves diligently, the risky project yields an output level of λA_t with probability q and zero otherwise. If instead the entrepreneur shirks, no output is produced while the entrepreneur obtains some private benefits $\beta\lambda A_t$ with probability q, with $\beta < 1$. This means that private benefits are just a fraction of the output that would be obtained in the case of success of the project, which implies that behaving diligently is socially optimal. Private benefits are measured in output units and they cannot be sized by external investors. The entrepreneur's behavior is not observable, so private benefits induce an agency problem. This limits entrepreneurs' ability to diversify business risk. The assumption that private benefits are obtained with probability q implies that shirking has no advantage in terms of risk relative to being diligent. We are also implicitly assuming that the safe project cannot generate any private benefit. As it will become clear below, the assumption is without loss of generality, provided that behaving diligently is socially optimal. Given this formulation, the expected return of a unit of capital invested in the risky project (if the entrepreneur behaves diligently) is equal to

$$\mu_r = q\lambda$$

while the *variance* of the project return is equal to

$$\sigma = \mu_r \lambda - \mu_r^2.$$

Increasing λ, while keeping μ_r fixed, implies an increase in the risk of the project for given expected return: as λ increases the success probability of the project falls but, in the case of success, its return is higher. So the parameter λ measures the *underlying idiosyncratic risk* in the sector: changes in λ have no consequences on the return of a well diversified portfolio, but they can influence the choices of an undiversified entrepreneur. A higher λ implies that a successful innovation is more valuable, but its probability of success is lower. This may be the result of fiercer competition in the markets served by the firm (say due to globalization), or by faster technological progress, that makes innovation more competitive.

Funds are provided by *investors* who are risk-neutral and discount future payments at an interest rate that for simplicity we normalize to zero. The individual supply is infinitesimal, but the aggregate number of investors is large enough to guarantee that the aggregate supply of funds is perfectly elastic at the given interest rate. This implies that financial markets are perfectly competitive and the equilibrium interest rate is zero. This could characterize an open economy with perfect capital mobility. Alternatively, one could think this corresponds to the equilibrium of our economy under autarky, since with a zero interest rate the capital market clears—i.e., the within period aggregate demand and aggregate supply of capital are both equal to A_t.

We also make the following two simplifying assumptions:

$$\mu_r > \mu_s \geq 1, \tag{8}$$

$$\underline{\tau} > 2\lambda. \tag{9}$$

Assumption 8 implies that, in the absence of financial frictions, operating the risky project would be socially optimal. Assumption 9 guarantees instead that the marginal utility of consumption is positive for any possible relevant value of consumption and propensity to take risk of entrepreneurs. Finally notice that the assumption that entrepreneurs have A_t unit of wealth and that a project involves A_t unit of investment implies that no entrepreneur is financially constrained. In the closed economy interpretation of the model, this assumption also guarantees that the capital market clears at the zero interest rate. So suboptimal investment decisions could result only from lack of risk diversification opportunities.

Business risk affects entrepreneurial activity and innovation, which are key determinants of productivity growth as in the Schumpeterian paradigm reviewed by Grossman & Helpman (1991) and Aghion & Howitt (1998). Similar specifications have been commonly used in the endogenous growth literature, at least since Romer (1990). We suppose that innovation is risky so that the rate of growth of aggregate sectoral technology γ_t is proportional to the number of successful risky projects with a factor of proportionality η per unit size of the innovation so that

$$\frac{A_{t+1} - A_t}{A_t} = \gamma_t = \eta \lambda q(1 - \rho_t) \tag{10}$$

where ρ_t denotes the fraction of entrepreneurs investing in the safe project. For simplicity here we are assuming that each successful risky project induces an intertemporal technological externality equal to $\eta\lambda > 0$ where λ is the size of the innovation, while safe projects produce no externality. The key assumption is that risky projects generate stronger technological spillovers than safer, more conservative projects.

A.2 The entrepreneur's problem

The entrepreneur must decide the type of project (risky or safe) and how to invest his wealth (whether in the project or in financial markets). To finance the project, the entrepreneur can sell equity in financial markets. Equity entitles external investors to a fraction $1 - \alpha$ of the revenue (if any) generated by the project. Selling equity allows the entrepreneur to fund a fraction $1 - i$ of the project investment with external finance. The entrepreneur can also reinvest the proceeds of selling shares in financial markets. This can guarantee the entrepreneur some income even if the project fails. One can also think of this income as a wage paid to the entrepreneur for managing the firm. Thus the combination of equity and reinvestment in financial markets allows the entrepreneur both to appropriate a fraction α of the cash flow generated by the project and a constant income θ per unit of capital invested in the project. The risk-free component of the project return θ reflects the insurance possibilities induced by institutional arrangements. Notice that, since any other wealth of the entrepreneur cannot be seized by external investors, θ has to be non-negative. The analysis below makes clear however that this constraint will never bind in equilibrium.

Once divided by A_t the time t expected consumption of the entrepreneur, conditional on the choice of the risky project ($j = r$) or the safe project ($j = s$), can be expressed as

$$E_j \left(c/A_t\right) = E\left[\alpha\tilde{\lambda} + \theta + (1 - i)\right] = \alpha\mu_j + \theta + (1 - i) \tag{11}$$

where $1 - i$ denotes the part of the project financed externally and $E_j(\tilde{\lambda}) \equiv \mu_j$. Analogously the second moment of the entrepreneur's consumption, once divided by A_t^2 is given by

$$E_j \left(c/A_t\right)^2 = E\left[\alpha\tilde{\lambda} + \theta + (1 - i)\right]^2 \tag{12}$$

which is again conditional on the type of project j chosen. Now notice that the participation constraint for financiers implies that

$$(1 - \alpha)\mu_j = \theta + (1 - i),$$

which says that the expected payments received by financiers must be equal to the present value of their disbursements. This constraint holds as an equality because of perfect competition in financial markets. Using this result to substitute for $\theta+(1-i)$ into (11) and (12) and after some algebra, we obtain that, if the safe project is chosen, the expected utility of consumption once divided by A_t is equal to

$$E_s\left[U_t(c)/A_t\right]=\mu_s-\frac{1}{\tau}\mu_s^2, \tag{13}$$

which is independent of α. For the risky project, an analogous substitution yields:

$$E_r\left[U_t(c)/A_t\right]=\mu_r-\frac{1}{\tau}\left[\mu_r^2+\alpha^2\mu_r\left(\lambda-\mu_r\right)\right]. \tag{14}$$

If the risky project is chosen, the problem of the entrepreneur can then be written as

$$\max_{\alpha} E_r\left[U_t(c)/A_t\right] \tag{15}$$

subject to

$$\alpha\geq\beta \tag{16}$$

where this last constraint is the incentive compatibility constraint for the entrepreneur, which imposes that the entrepreneur prefers behaving diligently to shirking. This expression is so simple because of the assumptions that private benefits are stochastic and measured in output units. To solve the problem note that (16) will always hold as an equality, since (14) implies that $E_r\left[U_t(c)/A_t\right]$ is strictly decreasing in α. Thus the equilibrium expected utility under the choice of a risky project is given by (14) with $\alpha=\beta$.

Now we can come back to the first stage of the entrepreneur's problem, which determines the choice of the project. Clearly the entrepreneur will choose to invest in the risky project if $E_s\left[U_t(c)\right]\leq E_r\left[U_t(c)\right]$, which after using (13) and (14) can be simplified to

$$(\mu_r-\mu_s)-\frac{1}{\tau}\left(\mu_r^2-\mu_s^2\right)\geq\frac{1}{\tau}\beta^2\sigma,$$

that is less likely to hold if either σ or β are high. From the previous expression we obtain a critical threshold

$$\tau^*=\frac{\beta^2\sigma+\left(\mu_r^2-\mu_s^2\right)}{\mu_r-\mu_s} \tag{17}$$

such that the entrepreneur will invest in the safe project only if his propensity to take risk is lower than τ^*. As a result the fraction of entrepreneurs investing in the safe project is given by

$$\rho=\max\left[0,\min\left(1,\frac{\tau^*-\underline{\tau}}{\overline{\tau}-\underline{\tau}}\right)\right], \tag{18}$$

which is constant and independent of time. Given (10), the constant over time productivity growth rate is equal to

$$\gamma=\eta\mu_r(1-\rho) \tag{19}$$

The average variability of projects returns depend on the idiosyncratic risk of risky projects and on the share of entrepreneurs that invest in them. The *observed* average idiosyncratic risk of project returns can therefore be expressed as

$$\omega = (1 - \rho)\, \sigma, \tag{20}$$

since just a fraction $(1 - \rho)$ of entrepreneurs invest in risky projects, each of them having idiosyncratic risk of returns σ.

A.3 The two main text implications

The previous model has two key empirical implications discussed as Proposition 1 and 2 in the main text. One is that the observed average idiosyncratic risk in the sector ω is endogenous to the risk diversification opportunities β and the level of underlying risk σ. Another is that the effect of idiosyncratic risk on the sectoral growth rate varies depending on the level of underlying idiosyncratic risk σ and risk diversification opportunities β.

Proposition 1 *In economies with high risk diversification opportunities (low β) the observed level of idiosyncratic risk ω accurately measures the underlying idiosyncratic risk σ. When risk diversification opportunities are low, observed risk is endogenous and moves less than one-for-one with underlying risk. The attenuation effect is larger the lower the risk diversification opportunities. In a regression of productivity growth on observed risk, the sign of the endogeneity bias can go either way, and it can be strong enough to lead to the erroneous conclusion that higher idiosyncratic risk improves economic performance.*

Proof. Assumption 9 guarantees that there exists a sufficiently low (yet positive) value of β such that τ^* in (17) is equal to $\underline{\tau}$, so that $\rho = 0$. For this (or any lower) value of β the observed idiosyncratic risk in the sector ω is equal to the underlying idiosyncratic risk σ. But when risk diversification opportunities are sufficiently low and idiosyncratic volatility high enough to make $\rho > 0$, ω becomes a generally (very) imperfect measure of σ. To see this, assume that $0 < \rho < 1$, then, after using (17), taking derivatives in (2) yields

$$\frac{\partial \omega}{\partial \sigma} = 1 - \rho - \frac{\beta^2 \sigma}{(\bar{\tau} - \underline{\tau})\,(\mu_r - \mu_s)} < 1. \tag{21}$$

Moreover, using (18) and (9), we can also see that

$$\lim_{\beta \to 0} \frac{\partial \omega}{\partial \sigma} = 1. \tag{22}$$

Equation (21) implies that for sufficiently large β the derivative $\partial\omega/\partial\sigma$ is strictly less than one, possibly negative, and, since ρ is decreasing in β, decreasing in the level of risk diversification opportunities, β. When β is low enough, $\partial\omega/\partial\sigma = 1$, which means that the observed level of idiosyncratic risk accurately measures the underlying idiosyncratic risk in the sector. The fact that the derivative of ω with respect to σ could be smaller than one implies an endogeneity bias of generally uncertain sign, when running a regression of γ on ω. To see this point more clearly assume that risk diversification opportunities are high (but not so high as to induce $\rho = 0$), so that $\partial\omega/\partial\sigma$ is positive and strictly less than one. In this case, a higher σ (due to an increase in λ) tends to lead to a fall in γ and to a less than a one-for-one increase in ω, so that an OLS estimate of the ω-coefficient tends to over-estimate the negative effects of an increase in idiosyncratic risk σ on γ. When instead risk diversification opportunities are so low that $\partial\omega/\partial\sigma$ turns negative, an increase in σ makes γ and ω both fall. In this case an OLS regression of γ on ω would yield a positive coefficient on the variable ω, which would misleadingly suggest that higher risk leads to higher productivity growth. ■

Proposition 2 instead says that:

Proposition 2 *An increase in underlying idiosyncratic risk σ reduces productivity growth γ. The effect is stronger the worse the risk diversification opportunities (larger β).*

Proof. Suppose that we are not at a corner solution so that $0 < \rho < 1$. Using (17) to substitute for τ^* in the expression for ρ in (1) yields

$$\gamma = \frac{\eta\mu_r\left(\bar{\tau} - \mu_r - \mu_s\right)}{\bar{\tau} - \underline{\tau}} - \frac{\eta\mu_r}{\left(\mu_r - \mu_s\right)\left(\bar{\tau} - \underline{\tau}\right)} \cdot \beta^2\sigma, \tag{23}$$

which says that, when σ increases, less entrepreneurs invest in the high-risk-high-return project, so that the productivity growth rate falls. The effect is stronger the less diversified the entrepreneurs are. When instead ρ is equal to zero or one, σ has marginally no effect on γ. ■

A.4 Biases in Rajan and Zingales regressions

Consider now the n independent sectors, each characterized by a different level of idiosyncratic business risk, denoted by σ_j, $j = 1, ..., n$. Sectoral differences in business risk might be due to differences in technology, in the degree of competition or in the riskiness of innovation activities. Entrepreneurs have project opportunities in one sector. Sectoral productivity growth depends on the number of successful risky projects within the sector, exactly as in (1). The key assumption is that an innovation induces stronger technological spillovers within the sector than across sectors. Now consider two countries that differ in risk diversification opportunities β. Proposition 2 predicts that the country with worse diversification opportunities will grow relatively less in sectors with greater idiosyncratic risk. We can therefore relate the growth performance of a sector within a country to the corresponding level of idiosyncratic risk in the sector and then check how the relationship differs for countries with different risk diversification opportunities. In terms of the model this amounts to checking how $\partial\gamma/\partial\sigma$ differs for countries with different β, which measures the sign and magnitude of the second order partial derivative $\frac{\partial^2\gamma}{\partial\beta\partial\sigma}$. Based on this logic, we test whether sectors with higher idiosyncratic risk perform relatively worse in countries with less risk diversification opportunities. The empirical challenge is that observed risk ω is endogenous, see Proposition 1. Failing to recognize this leads to an important endogeneity bias.

To emphasize the distinction between observed and underlying risk, we have focused the discussion on the sign and magnitude of the correlation between observed risk and growth. But as discussed, our empirical strategy is based on cross-country industry data. In this context, the bias will depend both on how observed and underlying risk are related—i.e the sign and magnitude of the $\partial\omega/\partial\sigma$ derivative—and on how underlying risk differs in countries with different risk diversification opportunities. To see the determinants of the bias more formally, we can use (23) and (21) to express the derivative of sectoral performance with respect to observed risk ω as equal to

$$\frac{\partial\gamma}{\partial\omega} \equiv \frac{\partial\gamma/\partial\sigma}{\partial\omega/\partial\sigma} = -\frac{(\mu_r - \mu_s)}{(\mu_r - \mu_s)(1-\rho)(\bar{\tau} - \underline{\tau})\beta^{-2} - \sigma}. \tag{24}$$

This corresponds to the OLS estimates of the a_1 coefficient in the regression analysis with cross-country industry data. It is easy to check that, if $\partial\omega/\partial\sigma > 0$, the denominator is positive and decreasing in β. But whether the above derivative will be higher or lower in countries with different risk diversification opportunities will now also depend on how β covaries with σ. For example, if underlying risk σ is sufficiently lower in countries with higher β, using observed risk could misleadingly lead to even reject the hypothesis that idiosyncratic risk has bigger negative effects on economic performance in countries with lower risk diversification opportunities—i.e., $\frac{\partial^2\gamma}{\partial\beta\partial\omega}$ could be found to be positive. Generally, a positive $\frac{\partial^2\gamma}{\partial\beta\partial\omega}$ derivative (which is equivalent to the OLS estimates of the a_1 coefficient in the regression analysis) and a negative $\frac{\partial^2\gamma}{\partial\beta\partial\sigma}$ derivative (which is equivalent to the IV estimates of the a_1 coefficient in the regression analysis) are due to the fact that observed risk is endogenous, and that countries with worse risk diversification opportunities happen to have lower underlying risk.