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DOES IDIOSYNCRATIC BUSINESS RISK MATTER?

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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 volatility. Given that volatility is endogenous with respect to risk diversification opportunities, we instrument its value at the country-sector level with the corresponding sectoral volatility in the US, a country where idiosyncratic business risk can be more easily diversified away. Diversification measures are instrumented using demographic changes induced by World War II. We also provide firm-level evidence suggesting that firms controlled by less diversified owners display lower mean and dispersion of productivity growth.

JEL classification: 04, F3, G1.
Key words: Entrepreneurship, risk diversification, growth.

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1 Introduction

In standard Arrow-Debreu economies with complete markets, idiosyncratic risk can be fully diversified away and it is irrelevant for equilibrium outcomes. But as emphasized by Townsend (1978) and Holmstrom (1979), among others, full risk diversification is costly and much theoretical research has analyzed how various financial frictions can prevent it, hampering aggregate productivity, output, and capital accumulation as in, for example, 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 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 growth. A key issue in identifying the effects of idiosyncratic risk is that the volatility of observed growth or of any other economic outcome (in brief observed risk) could be a very imperfect measure of the true underlying risk, which determines institutional arrangements and shapes agents' decisions. For example, principal agent models as in Holmstrom & Milgrom (1987) show that a trade off generally exists between risk-sharing benefits and provision of incentives and as a result observed risk only indirectly measures underlying risk. More recently, Aghion, Angeletos, Banerjee & Manova (2007), Thesmar & Thoenig (2000), and Thesmar & Thoenig (2004) have also stressed that observed risk is endogenous to the market structure and to the risk diversification opportunities available in the economy, since firms react to changes in the economic environment by modifying their organization structure and their innovation activities. Coles, Daniel & Naveen (2006), Fischer (2008) and Bartram, Brown & Stulz (2009) provides more direct evidence that observed idiosyncratic risk is endogenous and can be influenced by firms' decisions.

In this paper we provide evidence on the effects of idiosyncratic risk on growth. To analyze the issue we consider a simple extension of the moral hazard model by Holmstrom & Tirole (1997) where risk averse entrepreneurs can choose between projects with different risk-returns tradeoffs. To solve the ensuing agency problem (which is a source of financial frictions), entrepreneurs have to partly finance the business venture with their own wealth and so idiosyncratic business risk cannot be fully diversified away.[1] Because of this, en-

[1] 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-up new businesses is substantial.

trepreneurs may choose projects that are strictly dominated from the point of view of a well diversified portfolio, just because they have a lower idiosyncratic risk. This hinders innovation, entrepreneurial activity, and growth. The model delivers two key predictions that would be common to a vast class of models: first, the effects of idiosyncratic risk on growth strictly depend on the risk diversification opportunities in the economy—with zero effects for big enough diversification opportunities and negative and increasing effects as diversifying risk becomes sufficiently difficult; second, the observed volatility of the 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 risk 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 county-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 focus on publicly traded firms, because this is arguably a better measure of the exogenous level of idiosyncratic risk that would be observed in a perfectly diversified firm.[2] We then follow Campbell, Lettau, Malkiel & Xu (2001) and decompose the overall variability of returns into a market and a firm idiosyncratic component. Observed risk is measured by the volatility of the idiosyncratic component. As stressed above, a key problem in estimating the effects of risk on growth is that observed risk is endogenous with respect to the diversification opportunities in the economy. To tackle this issue, we instrument a country's sector-level idiosyncratic volatility with the analogue 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 as in the US. We allow the relation 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.

In our model differences in risk diversification are due to financial frictions but in practice several other factors, including cultural aspects or some historical events, can limit entrepreneurs' ability to diversify away the risk of their business. To characterize the degree of diversification of business risk in the country and to identify its causal effects, we

[2]See Pagano & Roell (1998) for an analysis of how the legal environment affects the decision of companies to go public.

then 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. In fact, as shown by Moskowitz & Vissing-Jørgensen (2002), owners of family firms tend to have a substantial share of their wealth invested in their business. Since risk diversification measures could be endogenous, we instrument their value with changes in the demographic structure of the country's population due to World War II, which we think as an exogenous shock to the possibility to transmit businesses from one generation to the next and thereby to the country's business ownership structure and to the degree of diversification of business risk.

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 substantial and stand clearly only after explicitly accounting for the fact that observed and underlying risk differ. We also find important country specific differences in idiosyncratic risk. Results 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.

We also use firm level data for Italy to analyze the effects of firm ownership on firm performance and observed risk. We distinguish between firms controlled by an individual or family on one side and by holdings, financial institutions or foreign entities on the other. The latter are likely to hold a diversified portfolio of assets and therefore to act as a fully diversified entity. We find support for the key theoretical prediction that more diversified firms perform better on average, but with a greater risk of experiencing particularly bad outcomes—as measured by the mass on the left tail of the distribution of firms productivity growth.

Other papers have also analyzed the effects of idiosyncratic risk on economic performance. Caggese (2006) provides micro evidence that firms with higher observed volatility of profits invest in safer innovation activities with the effect varying depending on the degree of diversification of the firm's shareholders. He does not however address the issue of the endogeneity of risk and of the firm's ownership structure. Castro, Clementi & MacDonald (2008) use the Rajan and Zingales' methodology to analyze the effect of idiosyncratic risk on sectoral employment size using cross country-sector data and imposing that underlying idiosyncratic risk in every country is the same as observed risk in the US. Cũnat & Melitz (2007) and Manova (2007) use a similar approach but focus on the effects of risk on trade rather than on sectoral size. These 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 important country specific differences in idiosyncratic risk and we find that the effects of risk on growth are magnified by taking them into account.

Our 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 countries, for which differences in technology and in products 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.

Our empirical approach extends the methodology by 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 the US data. Ciccone & Papaioannou (2007) extend the methodology to show how to remove error in the US measure that, if not properly taken into account, could bias estimates. 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.

Our findings can help explaining the diverging productivity performance of the US relative to other Continental European countries over the recent past. Several authors have argued that the degree of "turbulence" of the world economy has increased due to an acceleration in the pace of technological progress and to the increased globalization of product markets, see for example Ljungqvist & Sargent (1998). Indeed Campbell et al. (2001), Comin & Mulani (2006), and Comin & Philippon (2005) provide evidence that idiosyncratic business risk in the US is today greater than it was back in the 70's. Thesmar

& Thoenig (2000) document similar evidence for Europe.[3] But the effects of idiosyncratic risk on growth vary depending on the risk diversification opportunities in the economy and they are arguably most damaging to Europe, where greater financial frictions prevent entrepreneurs from diversifying risk. In accordance with this interpretation, we find that the fall in productivity growth since the increase in economic turbulence in the 70's has been more pronounced in less diversified economies.

The remainder of the paper is structured as follows. Section 2 presents the model. Section 3 describes the empirical methodology. Section 4 discusses the aggregate data. Section 5 presents results and performs several robustness exercises. Section 6 presents the firm level evidence. Section 7 concludes and discusses some implications for the recent widening in the US-Europe productivity gap.

2 The model

To analyze the effects of idiosyncratic risk in economies that differ in the level of risk diversification opportunities, we build on Holmstrom & Tirole (1997). The model is intended to highlight key issues in identifying the effects of idiosyncratic risk on growth, arguably common to a vast class of models.

2.1 Assumptions

The economy lasts one period. There is a measure one of entrepreneurs with an initial amount of wealth equal to one and quadratic consumption preferences:

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

Here τ denotes the propensity to take risk of entrepreneurs, which differs across entrepreneurs according to a uniform distribution with support $[\underline{\tau}, \bar{\tau}]$. Entrepreneurs can invest in a project that costs one unit of wealth. Projects could be risky or safe, with expected returns μ_r and $\mu_s < \mu_r$, respectively. Project choice is irreversible. The safe project yields μ_s with certainty while, if the entrepreneur behaves diligently, the risky project yields an output level of y with probability q and zero otherwise. If instead the entrepreneur shirks, no output is produced while the entrepreneur obtains some private benefits βy with probability q, with $\beta < 1$. This means that private benefits are just a fraction of the output that would be

[3] Recently, Davis, Haltiwanger, Jarmin & Miranda (2006) has shown that idiosyncratic observed volatility in the US has increased only among publicly traded firms, while it has decreased among private firms. This might be because private firms, typically less diversified, have responded to the increase in underlying idiosyncratic risk by focusing on safer more conservatives activities.

obtained in 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.[4] Given this formulation the expected return of the risky project (if the entrepreneur behaves diligently) is equal to

$$\mu_r = qy$$

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

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

Increasing y, while keeping μ_r fixed, implies an increase in the risk of the project for given expected return: as y increases the success probability of the project falls but, in case of success, its return is higher. So the parameter y measures the *underlying idiosyncratic risk* in the economy: changes in y have no consequences on the return of a well diversified portfolio, but they can influence the choices of an undiversified entrepreneur. A higher y 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.

We also make the following two simplifying assumptions:

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

$$\underline{\tau} > 2y. \tag{2}$$

Assumption 1 implies that, in the absence of financial frictions, operating the risky project would be socially optimal. Assumption 2 guarantees instead that the marginal utility of consumption is positive for any possible relevant value of consumption and propensity to

[4]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 (of course provided that behaving diligently is socially optimal).

take risk of entrepreneurs. Finally notice that the assumption that entrepreneurs have one unit of wealth and that a project involves one unit of investment implies that no entrepreneur is financially constrained. So suboptimal investment decisions could result just from lack of risk diversification opportunities.

2.2 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 part of the project investment $1-i$ with external finance. The entrepreneur can also reinvest the proceeds of the selling of shares in financial markets. This can guarantee the entrepreneur some income even if the project fails.[5] Thus the combination of equity and reinvestment in financial markets allow the entrepreneur both to appropriate a fraction α of the cash flow generated by the project and a constant income t per unit of capital invested in the project. The risk-free component of the project return, t, reflects the insurance possibilities induced by institutional arrangements.[6]

The 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

$$E_j(c) = E\left[\alpha\tilde{y} + t + (1-i)\right] = \alpha\mu_j + t + (1-i) \tag{3}$$

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

$$E_j(c^2) = E\left[\alpha\tilde{y} + t + (1-i)\right]^2 \tag{4}$$

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

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

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. After using this result to substitute for $t+(1-i)$ into (3)

[5] One can also think of this income as a wage paid to the entrepreneur for managing the firm.

[6] Notice that, since any other wealth of the entrepreneur cannot be seized by external investors, t has to be non-negative. The analysis below makes clear however that this constraint will never bind in equilibrium.

and (4) and after some algebra, we obtain that, if the safe project is chosen, the expected utility of consumption is equal to

$$E_s\left[U(c)\right] = \mu_s - \frac{1}{\tau}\mu_s^2, \tag{5}$$

which is independent of α. An analogous substitution, under the assumption that the risky project is chosen, yields that

$$E_r\left[U(c)\right] = \mu_r - \frac{1}{\tau}\left[\mu_r^2 + \alpha^2\mu_r\left(y - \mu_r\right)\right]. \tag{6}$$

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

$$\max_{\alpha} E_r\left[U(c)\right] \tag{7}$$

subject to

$$\alpha \geq \beta \tag{8}$$

where this last constrain 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 (8) will always hold as an equality, since (6) implies that $E_r\left[U(c)\right]$ is strictly decreasing in α. Thus the equilibrium expected utility under the choice of a risky project is given by (6) 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 $E_s\left[U(c)\right] \leq E_r\left[U(c)\right]$, which after using (5) and (6) can be simplified to

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

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

$$\tau^* = \frac{\beta^2\sigma + \left(\mu_r^2 - \mu_s^2\right)}{\mu_r - \mu_s} \tag{9}$$

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

$$\rho = \max\left[0, \min\left(1, \frac{\tau^* - \underline{\tau}}{\bar{\tau} - \underline{\tau}}\right)\right], \tag{10}$$

which allows to write the productivity of the economy at the end of the period as equal to

$$\gamma = \mu_r - \rho \left(\mu_r - \mu_s \right) \tag{11}$$

while the *observed* average idiosyncratic risk in the economy is given by

$$\omega = (1 - \rho) \, \sigma \tag{12}$$

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

2.3 Some implications

The previous model has two key empirical implications. One is that the effect of idiosyncratic risk on an economy's performance varies depending on the level of underlying idiosyncratic risk σ and risk diversification opportunities β. Another is that the observed average idiosyncratic risk in the economy ω is endogenous to the risk diversification opportunities β and the level of underlying risk σ.

To see the first implication more formally, suppose that we are not at a corner solution so that $0 < \rho < 1$. Using (9) to substitute for τ^* in the expression for ρ in (11) yields

$$\gamma = \mu_r + \frac{\underline{\tau} \left(\mu_r - \mu_s \right) - \left(\mu_r^2 - \mu_s^2 \right)}{\bar{\tau} - \underline{\tau}} - \frac{1}{\bar{\tau} - \underline{\tau}} \cdot \beta^2 \sigma, \tag{13}$$

which says that, when risk σ increases (say because y rises), less entrepreneurs invest in the high-risk-high-return project, so the productivity in the economy 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. We conclude that:

Implication 1 *An increase in the underlying idiosyncratic risk in the economy has a negative impact on the economy's productivity. The effect is stronger, the lower the opportunities to diversify risk.*

Assumption 2 guarantees that it exists a sufficiently low (yet positive) value of β such that τ^* in (9) is equal to $\underline{\tau}$, so that $\rho = 0$. For this (or any lower) value of β the observed idiosyncratic risk in the economy 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 (9), taking derivatives in (12) yields

$$\frac{\partial \omega}{\partial \sigma} = 1 - \rho - \frac{\beta^2 \sigma}{\left(\bar{\tau} - \underline{\tau} \right) \left(\mu_r - \mu_s \right)}. \tag{14}$$

Moreover, using (10) and (2), we can also see that

$$\lim_{\beta \to 0} \frac{\partial \omega}{\partial \sigma} = 1. \tag{15}$$

Equation (14) 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 economy. 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 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 y) tends to lead to a fall in γ and to a less than a one-for-one increase in ω. In this case, 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. In brief, we have that:

Implication 2 *In economies with high risk diversification opportunities the observed level of idiosyncratic risk accurately measures the underlying idiosyncratic risk in the economy. When risk diversification opportunities are low, the observed risk is endogenous and it is imperfectly related to underlying 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.*

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. In practice, our empirical strategy below is based on relating the performance of a sector within country to the corresponding level of idiosyncratic risk and then analyzing how the relation differs for countries with different risk diversification opportunities. In terms of the model this amounts to checking how the $\partial\omega/\partial\sigma$ derivative differs for countries with different β, which identifies the effects on γ of the interaction term between β and σ in equation (13). Again failing to recognize the distinction between observed and underlying risk can lead to important biases, that would now depend not only on how observed and underlying risk are related—i.e the sign and magnitude of the $\partial\omega/\partial\sigma$ derivative—but also on how underlying risk differs in countries with different risk diversification opportunities. To see the issue more formally, one can use (13) and (14) to express the derivative of sectoral performance with respect to observed risk ω as equal to

$$\frac{\partial\gamma}{\partial\omega} \equiv \frac{\partial\gamma/\partial\sigma}{\partial\omega/\partial\sigma} = -\frac{(\mu_r - \mu_s)}{(\mu_r - \mu_s)(1-\rho)(\bar{\tau} - \underline{\tau})\beta^{-2} - \sigma}. \tag{16}$$

It is easy to check that, if $\partial\omega/\partial\sigma > 0$ (as it will turn out to be the case in the data), 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 effect on economic performance in countries with lower risk diversification opportunities.

3 Empirical methodology

The model delivers two key equations that, together with the availability of instruments for the level of underlying risk, can be used to test the effects of risk on aggregate performance. Equation (13) suggests running the following regression:

$$\gamma_{ji} = a_0 + a_1\beta_i \cdot \sigma_{ji} + a_2' X_{ji} + u_{ji} \tag{17}$$

where γ_{ji} is performance 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 sector and country dummies. As in Rajan & Zingales (1998), the regression (17) 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 relation differ 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 relation between ω and σ. In particular, after using (14) and (15) to linearize (12) with respect to underlying risk σ around a given level of idiosyncratic risk $\bar{\sigma}$, we obtain:

$$\omega_{ji} = c_{0i} + (1 - c\beta_i)\sigma_{ji} + \epsilon_{ji} \tag{18}$$

with $c > 0$. Here ω_{ji} is a measure of the observed idiosyncratic risk in a given sector j and country i and c_{0i} is a (possibly) country-specific term. The constant term can vary by country, because countries with different risk-diversification opportunities have different level of observed risk at the reference level of underlying risk $\bar{\sigma}$ around which the approximation is computed. The positive c coefficient implies that an increase in underlying risk translates into a less than a one-for-one increase in observed risk, with smaller increases the lower the risk diversification opportunities (i.e. the higher is β). This is one of the key implications of equation (14). When risk diversification opportunities are high enough observed and underlying risk move one- for-one, see equation (15).[7] 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 in the United States risk diversification opportunities are so large that the level of observed idiosyncratic risk accurately measures underlying risk (at least for listed firms). 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:

$$\sigma_{ji} = b_{0i} + \sum_{k=1}^{K} b_{ki} \left(\sigma_{jU}\right)^k + v_{ji} \tag{19}$$

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. The specifications allows for a non linear relation 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 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 the regression (17) and estimate the effects of risk on the economy 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

[7] Notice that we could generalize equation (18) by adding higher order terms in β_i.

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 (17) we can then use a Two-Stage-Least Square procedure: we can estimate the coefficients b_{ki}'s, then replace σ_{ji} in (17) with $\sum_{k=1}^{K} b_{ki} (\sigma_{jU})^k$ and finally estimate equation (19) by OLS.[8] To estimate the coefficients b_{ki}'s we can use (19) to replace σ_{ji} in (18). This yields a regression model in terms of observables that can be estimated by Non Linear Least Square. To identify the coefficient c we need some independent variation in the level of risk diversification opportunities for countries that share the same relation to risk in 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.[9]

A possible additional concern is that risk diversification measures could be endogenous to country specific differences in sectoral performance: for example because, as differences in sectoral performance increase, the incentive to diversify risk across sectors also increases, thereby leading to a change in the country's business ownership structure. While this is a somewhat less serious concern than the possible endogeneity with respect to aggregate growth, it is still an important one especially because our diversification measures are calculated at a date sometimes contemporaneous (or even posterior) to the period of reference of the sectoral performance measure. 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. To address these issues, we will also instrument diversification measures 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 today sectoral performance of countries.

[8]Notice that if we were to replace σ_{ji} in (17) 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 (17).

[9]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.

4 Data

We discuss the measure of idiosyncratic risk, sectoral performance, risk diversification and the instruments used.

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. 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, an industry, and a firm component. As a measure of fully diversifiable risk in perfect financial markets, we use the sum of the industry and firm volatility component. This will be our indicator of observed idiosyncratic risk at the level of year-country-sector.[10] Volatility is computed for 36 sectors, that closely follow the sectoral classification of Fama & French (1997). Appendix A describes in more detail the construction of the volatility measure and the sectoral classification used.

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 relevance of instruments we then 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 between the two measures, that can be appropriately exploited to improve estimation efficiency.

Table I 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). 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 I), although in most countries we have data for at least 20 sectors. A disturbing exception is however New

[10] We checked that the results are little affected by measuring idiosyncratic risk using just the firm volatility component rather than the sum of the industry and firm component. We do not report the results with this alternative volatility measure just to save space.

Zealand for which only three sectors are available.

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. The idea is that 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).[11] Productivity data are from the STAN database, compiled by the OECD. 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 II). 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 4 and 6 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 that we will use is taken from Bartelsmann, Scarpetta & Schivardi (2005), see their paper for details. Table II 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 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 difference weighting schemes to reduce the potential effects of measurement error.

[11] In terms of the model we think that future (logged) aggregate productivity y_n' is equal to the sum of past aggregate productivity and entrepreneurial success γ: $y_n' = y_n + \gamma$. Similar specifications have been commonly used in the endogenous growth literature, at least since Romer (1990).

[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. In fact, the regression analysis will use country dummies to control for cross country differences in average growth.

4.3 Diversification measures and their instruments

In our model differences in risk diversification are due only to financial frictions. In reality, several other factors, including cultural aspects or some exogenous shocks (see below), can limit entrepreneurs' ability to diversify away the risk of their business. 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 ownership 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.

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 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.[14] Due to data constraints,

[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. In any case our instrumental variable approach will take care of this concern, see below.

[14]In our model all entrepreneurs in a country own the same share of equity capital. To better see the mapping between this statistic and the theoretical framework, one could consider an extended version of the

these papers focus just on listed companies. Our assumption is that their ownership structure summarizes economy-wide properties. This is reasonable because listed firms account for a substantial part of the overall firms' capital and because the ownership structure of private and listed companies is influenced by common country-specific factors. We follow the methodology in Mueller & Philippon (2006) to improve the comparability of data across countries; some more details are discussed in Appendix B.

The instruments for risk diversification opportunities are based on World War II demographic changes in terms of military, civil, and Jews 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 Jews casualties because they are available and because they might characterize differential effects of the War on the country's demographic structure. 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 possibilities 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. We believe however that the opportunities to diversify business risk are best identified by direct measures of the businesses ownership structure in the country, for which the previous instruments seem also most appropriate to identify causal effects.

Table III 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 the three indicators of private benefits of controls, consistent

model where β is firm specific and it is a random draw from a given distribution whose mean is affected by the institutional environment. Then, the fraction of entrepreneurs who own more than a certain threshold of equity measures the average β in the distribution and it can be used to characterize the institutional environment.

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

with the notion that agency problems are an obstacle to risk diversification.

5 Results

We start by discussing the results of the first stage of the estimation procedure, that yields a characterization of the relation between idiosyncratic risk in the US and in other countries. We then turn to the estimate of 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 of 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's parameters in equation (19)—that characterizes the relation between underlying risk in the US and in other countries—and of the c parameter in equation (18)—that characterizes the relation 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 relation between US risk and country specific risk in equation (19) 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. The details of the estimation procedure are reported in Appendix C.

Table IV present the results. The estimated value for c is strictly 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's parameters in (19) are statistically significant for at least one of the two volatility indicators in 14 out of 20 cases. Just for 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 I. 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 relation 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 differ across countries, since a test strongly reject the null that the b's coefficients are equal across countries. This would go 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 (17). We start 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.[16] 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 V 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 at the end of Section 2, 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 V reports the results when using the US volatility measure by Campbell et al. (2001) based on CRSP data. Column 3 deals with the analogous 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

[16]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.

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 (that corresponds to the 25^{th} percentile of the family ownership distribution by country) than in Italy (that 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 level, 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.[17]

The previous estimation hinges on assuming that sectoral underlying risk is the same across countries. Column 4 in Table V 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 coherent with the idea that countries with less diversified ownership structure 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.

5.3 Robustness checks

To account for the endogeneity of family ownership, we instrument its interaction with volatility with demographic changes induced by WWII, also interacted. The first-stage regression (unreported for brevity) 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

[17] 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.

highly significant. A possible interpretation is that civilian casualties also characterize the destructive effects of WWII on the production system: with more destruction leading to more creation of family controlled firms after the war. Higher casualties among military and Jews may instead indicate a more dramatic reduction in the relatively younger portion of the population, which has made more difficult the transmission of family businesses from the War generation to the next. 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.

Column 1 in Table VI presents the results of the IV regressions. The interaction coefficient now falls to minus 14.6. The substantial 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 18th and 19th centuries, as constructed by Crouch (1993), are strong predictors of the today importance of family ownership in the country. 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 Machineries 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,[18] in column 2 of Table VI we present the results when weighting the observations with the country specific sectoral employment size. The coefficient drops (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 (2007) and Castro et al. (2008)—or

[18]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.

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 use just a cross section, taking the average of both productivity growth and volatility over a potentially long period of time.[19] 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 VI, where we compute average productivity and volatility separately for each of the six five-year periods that go 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.[20]

So far we have excluded New Zealand and Portugal from the sample, since their first stage results were somewhat dubious (see Table IV). Column 5 of Table VI 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 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

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

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

financial development proposed by Rajan & Zingales (1998).[21] The correlation coefficient between external dependence and our measure of idiosyncratic risk is .24; that 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— is -.42 and -.23, respectively. In the last two columns of Table VI we report the results of the basic regression when including also 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 also for other reasons than risk diversification. For example, Caselli & Gennaioli (2005) argue that the ownership structure affects the selection of managerial talent. Table VII 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 rule of law, anti-director rights and accounting standards.[22] 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 the indicator indicates 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 standard turning out not to be statistically significant at the 10% level. We

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

[22] To simplify comparison of results, the only difference relative to the previous tables is that the interaction term in equation (17) is computed using these alterative measures.

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 second two capture the 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 VIII. Odd columns present the un-weighted regressions, the even columns presents results with US weights. Results with alternative weighting schemes are similar and are not reported 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 fall substantially when considering these alternative measures, reducing the precision of the estimates.

6 Firm level evidence

We now explore some implications of the model using micro data. The data come from INVIND, a survey of medium to large size Italian manufacturing firms conducted every year by the Bank of Italy since 1982. The survey is representative of the population of manufacturing firms with at least 50 employees and collects several information on firms, including employment, sales, and the identity of the controlling shareholder (see Iranzo, Schivardi & Tosetti (2008) for a detailed description of the dataset). In particular, it distinguishes between companies owned by i) individuals and families, ii) holdings, iii) financial institutions, iv) foreign owners, and v) public entities. We exclude publicly owned firms from the analysis, as their objectives often differ from simple profit maximization. As before, we assume that owners of individual and family controlled firms are less diversified than other owners. In fact, holdings and financial institutions are likely to have a diversified portfolio of assets, while foreign firms are typically owned by multinationals or financial institutions, such as private equity funds, arguably diversified.

As before, we focus the analysis on productivity growth, measured by changes in sales per worker (data on value added are not available). To obtain a symmetric distribution of growth rates, that makes the graphical analysis below easier to interpret, growth is defined as follows:

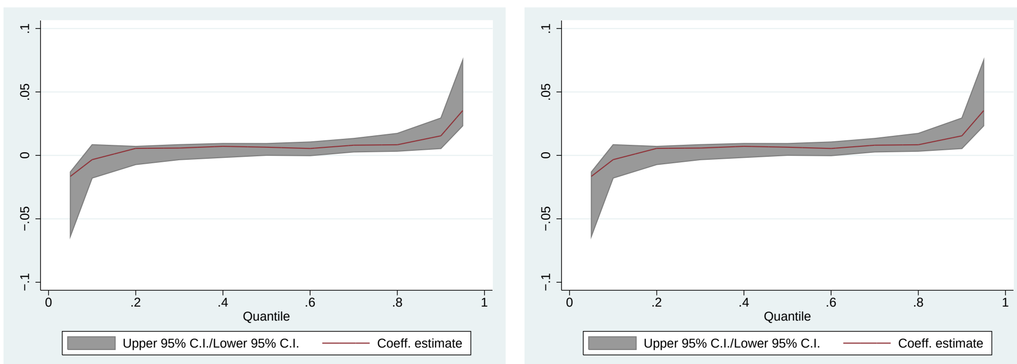
$$\Delta prod_t = \frac{prod_t - prod_{t-1}}{\frac{1}{2}(prod_t + prod_{t-1})}$$

which ranges between -2 (when time t productivity goes to zero or time $t-1$ productivity goes to infinity) and 2. Our prediction is that firms controlled by diversified owners on average grow more, but that their productivity performance is riskier. In particular, we expect more mass on the left tail of the distribution of productivity growth, which would imply a greater probability of experiencing particularly bad productivity outcomes. To test the hypothesis we start running a regression of productivity growth on a full set of year dummies to account for aggregate factors and then use the residual as our first measure of firm performance. We report the distribution of the residuals in the first two columns of Table IX. The results indicate that average growth is greater for diversified firms, although their productivity performance is more volatile. In particular, the distribution of productivity growth for non-family firms has a fatter left tail and the two distributions cross approximately at the 10^{th} percentile.

One could argue that the two groups of firms should be made more homogeneous. On one hand family firms might be more prominent in low-tech, low-risk sectors just for historical reasons. On the other hand, sectoral choice is endogenous since family firms might decide to operate in some specific sectors exactly because of their lower risk. Size and age might also differ systematically between the two groups, and it is well-know that small and young firms tend to exhibit more volatile growth rates (Bartelsmann et al. 2005). We therefore also experiment with a specification where we include seven sector dummies, four region dummies, firm age, firm employment size and finally lagged growth rate, to account for serial correlations in productivity growth. We report the percentiles of the distribution of the residuals in the last two columns of Table IX. The results are similar to those obtained with only year dummies as controls, although now the two distribution cross at the 25^{th}, which is substantially greater than before. This might be due to the fact that family firms are on average smaller and therefore their residual volatility decreases when taking this into account.

To go beyond descriptive evidence and check statistical significance we also run quantile regressions on a dummy equal to 1 for non family firms. The coefficient measures the difference in the two distributions at the specified quantile. The estimated coefficients, with the corresponding 5% confidence interval, for the case with only year dummies as control are reported in panel (a) of Figure 1. For most quantiles, the difference is positive and around 1%. But, in line with the descriptive evidence, we also find that for quantiles below the 20^{th} percentile the coefficient is negative and statistically significant.[23] In panel (b) we report the results when adding the previously discussed additional controls. Results change little, although differences become more pronounced. All in all, this evidence is consistent with the theoretical prediction that diversified firms tend to outperform family firms, but at the same time they face a significant larger chance of a substantially worse performance.

Figure 1: Coefficient of the dummies for diversified firms in the quantile regressions



(a) No controls

(b) Controls

7 Conclusions

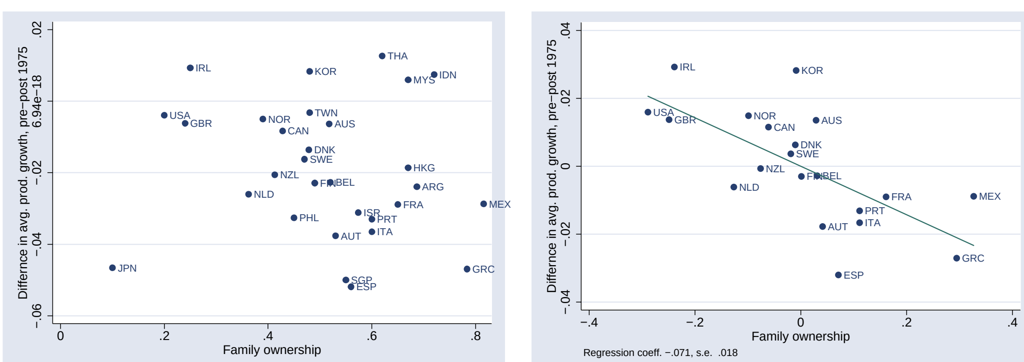
We used measures of diversification opportunities based on the ownership structure of businesses and we analyzed the effects of idiosyncratic business risk on sectoral performance in OECD countries. Since observed volatility is endogenous, we instrumented a country volatility with the corresponding volatility in the US, a country where entrepreneurs have better opportunities of diversifying away the idiosyncratic risk of their business. We found that countries with low levels of diversification perform relatively worse in sectors characterized by high idiosyncratic volatility. The implied effects are substantial and stand clearly only after taking into account that observed and underlying risk differ. The effects are magnified by allowing risk to vary by country and by instrumenting risk diversification opportunities with demographic changes induced by WWII. Results hold true with alternative measures

[23]We have excluded from the graph the first and last percentile, as their large values (in absolute terms) make the graph less readable. In fact, the coefficient is -.18 at the first percentile and .26 at the 99^{th}, both statistically significant, indicating large differences in the tails.

of growth performance (in terms of labor productivity, capital, value added, and business creation) or by using more direct measures of financial frictions that prevent entrepreneurs from diversifying away business risk.

The analysis has some interesting implications for the productivity growth experience of the US and Europe over the recent past. As discussed in the Introduction, there is evidence that idiosyncratic business risk has increased since the mid seventies, at least among OECD countries. This is sometimes attributed to the increased globalization of the markets served by companies or to the acceleration in the pace of technological progress documented by Gordon (2000), Greenwood & Yorokoglu (1997), and Jorgenson & Stiroh (2000). But the effects of idiosyncratic risk on growth are arguably more damaging to Europe, where greater financial frictions prevent entrepreneurs from diversifying risk. This may have discouraged European entrepreneurs from investing in new technologies and starting up new ventures, thereby causing a fall in productivity growth relative to the US. Figure 2 contains some suggestive evidence. In panel (a) we plot the difference in average productivity growth, computed from the Penn World Tables (Heston, Summers & Aten 2006), before and after 1975, against the share of of family controlled firms in the economy. The figure characterizes

Figure 2: Productivity growth changes, pre-post 1975 vs. family ownership



(a) All countries

(b) OECD countries, excluding Japan

a negative relation between changes in productivity growth and the importance of family ownership. Japan is the only clear outlier: it has the lowest share of family ownership and it has experienced one of the sharpest drop in productivity growth in the sample. Since Japan went through a major depression since the late eighties, whose main causes are possibly unrelated to idiosyncratic risk and ownership structure, in panel (b) of Figure 2 we restrict the sample to the group of OECD countries and we exclude Japan. The data line up along a negative line. Spain, Italy and Germany has done quite poorly in terms of productivity growth, while the US, Great Britain, and Ireland have experienced a increase in productivity growth. These last are also the countries with the lowest shares of family firms in the economy. Although just suggestive, this hints at a potentially interesting link. Our analysis provides a causal interpretation for this evidence and it suggests that the interaction of the increase in idiosyncratic risk with differences in risk diversification opportunities across countries can explain a substantial part of the US-Europe gap in productivity growth emerged over the last two decades, as well as the important differences emerged across European countries. Investigating this issue further is an interesting 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 in 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 and 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 and 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 representativeness 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, fraction of value of top 20 firms controlled by families and 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, 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 (19) to the case where both measures are used. To maximize degree of freedom, we also exploit time series variation. Equation (19) then becomes:

$$\sigma_{jit} = b_{0i} + \sum_{k=1}^{3}\sum_{z=1}^{2} b_{kiz} \left(\sigma_{jUzt}\right)^k + v_{jit} \tag{20}$$

where $z = 1$ indicates Thompson Datastream and $z = 2$ CRSP. Using (20) to substitute for σ in equation (18), we obtain:

$$\omega_{jit} = d_i + \sum_{k=1}^{3}\sum_{z=1}^{2} b_{kiz}(1 - c\beta_i) \left(\sigma_{jUzt}\right)^k + \eta_{jit} \tag{21}$$

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 relation 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 (21), 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. We restrict the search for c over the range zero to two and a half. This is reasonable since negative values of c are excluded by assumption: a higher β must be associated with a lower observed volatility with respect to the underlying one. A value of c greater than two and a half would instead imply that more than three-quarters of the countries in the sample are on the negative side of the underlying-observed risk relation, which may be regarded as highly implausible. The standard errors, in Table

IV are those of the linear estimation procedure; the standard error for c is instead calculated by bootstrapping. Finally, many countries miss observations on some of the 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.

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Table I: 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 raw reports the volatility computed by Campbell et al. (2001). See Subsection 4.1 and Appendix B for sources and definitions.

Table II: 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 V and VI. 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 III: 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 IV: First stage regression

	Campbell		Datastream		$H_0 : b_1 = b_2 = 0$
	b_1	S.E.	b_2	S.E.	p-value
AUS	0.42	0.48	1.12^b	0.45	.009
CAN	0.92^c	0.52	0.73	0.59	.016
GRC	1.97^b	0.77	-1.09	1.47	.038
ITA	0.53	0.5	0.62	0.48	.116
JPN	0.58	0.45	0.73^c	0.39	.020
KOR	0.07	0.55	0.58	0.68	.636
MEX	0.52	0.84	3.39	2.63	.366
NZL	-0.02	10.44	-1.48	71.29	.998
PRT	10.21^a	1.11	-14.33^a	1.97	.000
AUT	0.82	0.85	-0.02	1.09	.447
BEL	0.69	0.51	0.76	0.47	.034
ESP	1.06^c	0.58	1.38^c	0.84	.008
FRA	1.02^b	0.51	1.44^a	0.53	.000
GBR	0.98^b	0.46	0.84^b	0.4	.001
GER	0.65	0.5	0.86^c	0.47	.020
NLD	0.85^c	0.48	0.9^b	0.43	.003
SCAND.	1.00^b	0.49	1.29^a	0.44	.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 (18) and (19) 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, see Appendix C for details.

Table V: Productivity growth, diversification and idiosyncratic risk

	Volatility measure			
	Observed	US-CRSP	US-Datastr.	Underlying
	[1]	[2]	[3]	[4]
Family$*$	2.84^c	-7.98^a	-6.47	-12.17^a
volatility	(1.55)	(2.84)	(5.71)	(5.12)
Initial prod.	$-.020^b$	$-.018^a$	$-.019^a$	$-.020^a$
	(.009)	(.006)	(.006)	(.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 VI: Robustness checks

	[1]	[2]	[3]	[4]	[5]	[6]	[7]
Family∗	-14.65^b	-8.55^a	-8.52^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	
Market Cap.						(.031)	
Ext.Dep∗							
Private Credit							-.054
							(.044)
Initial prod.	-.020^a	-.013^b	-.008^b	-.019^a	-.020^a	-.027^a	-.027^a
	(.006)	(.005)	(.004)	(.006)	(.006)	(.008)	(.008)
R^2	0.41	0.55	0.64	0.34	0.42	.42	.43
N. obs.	387	387	339	1010	395	213	213
LR stat. (p)	0.00						
Sargan (p)	0.44						
Weight	NO	OWN	US	NO	NO	NO	NO
IV	YES	NO	NO	NO	NO	NO	NO
Repetead CS	NO	NO	NO	YES	NO	NO	NO

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. In the IV regression of column [1] the interaction is instrumented with WWII casualties (see the main text for more details). 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 V keeping the observations for Portugal and New Zealand in the sample. In columns [6] and [7] we add the interaction terms proposed by Rajan & Zingales (1998) to identify the effects of financial development on growth. 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 VII: Other measures of diversification

	Widely held firms	Rule of law	Anti directors rights	Accounting standards
	[1]	[2]	[3]	[4]
Diversif.∗ volatility	12.87^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 VIII: 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	-.0-.018^c	-.0-.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%.

Table IX: Distribution of productivity growth for family and non family firms

	No controls		Controls	
	Family	Other	Family	Other
mean	-.006	.004	-.003	.002
s.d.	.223	.262	.188	.224
1^{st} percentile	-.599	-.781	-.533	-.710
5^{th} percentile	-.272	-.290	-.270	-.308
10^{th} percentile	-.191	-.184	-.184	-.188
25^{th} percentile	-.088	-.078	-.081	-.078
50^{th} percentile	-.006	.004	.001	.006
75^{th} percentile	.078	.086	.075	.084
90^{th} percentile	.161	.178	.174	.192
95^{th} percentile	.246	.285	.262	.311
99^{th} percentile	.603	.867	.592	.664

The table reports mean, standard deviation and the percentiles of the distribution of productivity growth using the INVIND data. Productivity growth is measured by the residuals of the regression of sales per worker on year dummies in the "No controls" columns and on year, sector, and geographical area dummies plus firm age, size and lagged sales per worker for the "Controls" columns. "Family" are family controlled firms and "Other" are firms controlled by other entities, see Section 6 in the main text for details.