Credit Supply and Corporate Innovations

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May 16, 2012

Abstract

We present causal evidence that financial development plays a key role in the technological progress. We focus on firms' innovative performance, measured by patent-based metrics, and exploit the staggered deregulation of banking activities across U.S. states during the 1980s and 1990s as a source of variation in the supply of credit. We find that the interstate banking deregulations had significant beneficial effects on innovation activities, especially for firms operating in sectors highly dependent upon external capital. This effect, which does not become evident until some years after deregulations, is partly driven by a greater ability of deregulated banks to diversify credit risk, and thus by a relaxation of financial constraints for bank-dependent firms.

JEL: G20, G30, O30

Keywords: financial development, banking deregulations, innovation, patents

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For insightful comments and suggestions we thank Morten Bennedsen, Bruno Cassiman, Andrea Fosfuri, Denis Gromb, Jens Iversen, Paloma Lopez-Garcia, Andrea Mina, Charlotte Ostergaard, Jesper Rangvid, Thomas Rønde, Ben Sand, Amit Seru, Battista Severgnini, as well as participants at the Zvi Griliches Research Workshop in Economics of Innovation (Barcelona), Druid Winter Conference (Cambridge), Nordic Finance Network Workshop (Oslo) and participants at the INSEAD brownbag seminar. Contact details:
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1. Introduction

Schumpeter argued that developed and well-functioning financial systems are essential for promoting innovation and long-term economic growth. As discussed by Jarayatne and Strahan (1996), this relationship arises via two possible mechanisms. First, a pure volume effect results when financial intermediaries channel savings to investment (Bencivenga and Smith 1991). Second, financial systems can increase the productivity of that investment by allocating funds to the most qualified firms (Greenwood and Jovanovic 1990; King and Levine 1993a). Our contribution is to establish the causal effect of financial development on firms' innovative performance.

Recent works document a positive association between innovation and various financing sources, such as equity issues (Atanassov et al. 2007; Brown et al. 2009), venture capital (Kortum and Lerner 2000), bank credit (Ayyagari et al. 2011; Benfratello et al. 2008; Nanda and Nicholas 2011; Winston Smith 2011), internal resources (Himmelberg and Petersen 1994) and IPOs (Bernstein 2012). As shown in Figure 1, our data also point to a positive correlation between patenting activity and loan supply in the U.S. While this evidence suggests that a wider access to external finance may favor innovation, the empirical study of finance's effects on innovation is plagued by the endogeneity of financial development. Arguably, general economic conditions, industry characteristics and other unobserved factors may influence both firms' innovation activities and credit availability. In addition, firms with higher value-added projects may have easier access to financing.

To overcome these concerns, we employ the staggered passage of interstate banking deregulations in the U.S. banking industry during the 1980s and 1990s to construct time variations in banking expansion across states. As documented in the banking literature, deregulations improved the quality of financial intermediation at large, e.g. by increasing credit supply, and encouraging the adoption of screening and monitoring technologies.

Importantly to us, a specific consequence of interstate banking deregulations was that by allowing banks to operate in different states they facilitated banks' geographic diversification of credit risk (Demyanyk et al. 2007; Goetz et al. 2011), which may in turn have encouraged lending to riskier borrowers, such as innovative firms.

Our main result indicates that interstate banking deregulations caused an increase in the innovative performance of manufacturing firms, as measured by patent-based metrics. We also find an increase in the *relevance* of the patenting activity, measured by citations received from future patents, in patenting *risk*, as well as in patenting *originality* and *generality*, suggesting that firms exposed to banking deregulations changed the technological nature of their research projects. These effects remain significant after controlling for common patenting trends, firm fixed effects and other confounding factors. In particular, controlling for the stock of research and development (R&D) we observe that following deregulations firms made a more productive use of their existing innovation inputs. Our results are also robust to controlling for other confounding policy changes, e.g. the deregulation of branching activities, which affected the competitiveness of the banking sector, and the passage of anti-takeover legislations, which changed the quality of state-level corporate governance.

Focusing on the supply side, we claim that the main channel behind our findings relates to a greater willingness of deregulated banks to take risk once they become more diversified geographically. Out-of-state banks may be thus willing to lend at more favorable terms – all the more so if credit risk in the deregulating state is less correlated with their existing exposure. We find that most of the increase in patenting activity occurred in states whose economies exhibited least comovement with the overall U.S. economy. Moreover, the effect on innovation was highest in those states where new out-of-state banks were entering from the states least comoving with that state. Also, the increase in patenting occurred in states that after deregulations experienced the largest change in the average geographic diversification of banks operating in that state.

While changes in credit supply can affect financing and investment decisions of a wide array of firms (Lemmon and Roberts 2010), we expect our results to vary depending on borrowers' characteristics, such as firms' reliance on external capital and their ability to access non-bank sources of funds. Consistent with this argument, we find that banking deregulations mostly affected the innovation policy of firms operating in industries highly dependent upon external capital, firms that tend to rely more on bank debt, i.e. with worse access to other segments of the capital market (Leary 2009), and firms that are young and informationally opaque (see e.g. Johnson 1997; Hadlock and James 2002). Furthermore, the effect of deregulations was prevalent among firms that experienced a high level of R&D expenditures following deregulations.

Existing studies have investigated the effect of banking deregulations on entrepreneurship and Schumpeterian creative destruction (Bertrand et al. 2007; Black and Strahan 2002; Cetorelli and Strahan 2006; Kerr and Nanda 2009, 2010).[1] Our findings add to this evidence by showing that interstate banking deregulations had strong effects on firms' innovative activities. Furthermore, we are able to link the two typical Schumpeterian effects of finance, i.e. the one on creative destruction and the one on innovation; we show that industries that experienced the highest increase in innovation following deregulations experienced a subsequent high increase in output.

We also contribute to a broad research on the relationship between financial development and economic growth (King and Levine 1993b; Jayaratne and Strahan 1996; Demirguc-Kunt and Maksimovic 1998).[2] Some determinants of this relationship, such as

[1] Existing works have also tested the effect of banking deregulations on income distribution (Beck et al. 2010), industry reallocation (Acharya et al. 2011; Bertrand et al. 2007) and trade flows (Michalski and Ors 2011; 2012). Recent works have also identified how deregulations impacted on the volatility (Correa and Suarez 2009) and cash holdings (Francis et al. 2011) of U.S. publicly listed firms.

[2] See Levine (2005) for a comprehensive review of this literature.

firm entry and entrepreneurial activity, have been widely analyzed (Black and Strahan 2002; Guiso et al. 2004; Cetorelli and Strahan 2006; Kerr and Nanda 2009). By contrast, direct evidence on the role of technological progress is still limited. In a closely related work, Benfratello et al. (2008) show that banking development in Italy increased the probability of process innovation. We extend this work in two ways. First, we adopt policy changes in the U.S. banking industry as a natural experiment to establish a *causal* effect on firms' innovation activities. Second, by focusing on corporate patenting we are able to establish an effect not only on the volume but also on the *technological nature*, *riskiness* and *relevance* of the innovation activities pursued.

Finally, our study complements a recent literature that, motivated by the recent financial crises and dry-up of credit, investigates how variations in the access to external finance affect corporate policies (Campello et al. 2010; Duchin et al. 2010; Leary 2009; Lemmon and Roberts 2010). Although we are unable to test whether the increase in innovation stems directly from bank lending, or indirectly, i.e. from non-bank institutions or investors that in turn benefited from deregulated banks, our results reinforce the notion that changes in the supply of credit have strong effects on corporate policies.

Section 2 describes the policy changes that transpired in the U.S. banking industry. Section 3 presents the data. Section 4 outlines our empirical methodology. Section 5 presents our main finding that interstate banking deregulations increased the number and quality of firm innovation. We then provide evidence on the channel explaining the increase in patenting: the improved ability of banks to diversity credit risk thus lending to riskier borrower (Section 6). We further show how our findings vary depending on firm and industry reliance on external finance (Section 7), and we discuss the association between innovation and output reallocation within industries (Section 8). Section 9 concludes.

2. Deregulations in the U.S. banking industry

Our identification strategy exploits the staggered passage of interstate banking deregulations[3], which represented a significant credit shock and, in particular, allowed banks to geographically diversify credit risk.

The geographic expansion of banking activities in the U.S. has been historically restricted by laws such as the McFadden Act of 1927 and the Douglas Amendment to the Bank Holding Company Act of 1956. However, during the period of 1970-90s U.S. states largely terminated the restrictions on the expansion of banks across and within states. The first state passing an interstate banking deregulation was Maine in 1978, followed by Alaska and New York in 1982. The wave of deregulations continued until the mid-1990s, when piecemeal changes in legislation, outside events, and competition among regulators led states to permit some type of interstate banking on a reciprocal or nonreciprocal basis (Johnson and Rice 2008). Table 1, Panel A, illustrates the timeline of interstate banking deregulations by state and year.

After interstate banking deregulations, bank holding companies operating in other states were allowed to acquire banks chartered in that state. Indeed, interstate deregulations let to an expansion of banks across state borders. In an average state, the fraction of assets from out-of-state bank holding companies rose from 0% in mid-1970 to 23% in mid-1990. The total number of banks fell from the mid-1970s to the mid-1990s, but this reduction was mostly driven by a drop in the number of small local banks (Kerr and Nanda 2009). Furthermore, the expansion of banks across states led to an increase in loan supply. Using state-level data on commercial bank loans provided by the Federal Deposit Insurance Corporation (FDIC) for the period of 1976-95, we find that, after controlling for year and

[3] The means of cross-state geographic expansion are: acquiring or establishing a charter in a state outside the main bank's home state (interstate *banking*) as well as acquiring or establishing a branch office, an office which is not separately chartered or capitalized, in a state outside the main bank's home state (interstate *branching*). Likewise, intrastate banking and branching refer to respective means of expansion within the main bank's home state (Johnson and Rice 2008).

state fixed effects, interstate banking deregulations were associated with a 8% increase in total net loan supply. Moreover, expanding out-of-state banks were endowed with more sophisticated monitoring technologies than local non-expanding banks (Dick and Lehnert 2010). Finally, lower restriction to banking across states improved the scope for geographic diversification (Goetz et al. 2011), allowing banks to finance risky projects, such as innovation activities, more freely and without increasing their overall risk.

Overall, the deregulation of the U.S. banking system was associated with several major improvements in the credit markets which set the stage for the emergence of "expansion-minded banks" (Black and Strahan 2002). Jayaratne and Strahan (1998) document that these deregulations were also associated with reduced loan losses and operating costs. These reduced costs were passed along to bank customers in the form of lower loan rates and are thought to have produced benefits for a broad array of economic activities (Kerr and Nanda 2009; 2010; Francis et al. 2011). The goal of our paper is to test whether interstate banking deregulations had a positive effect on firms' innovative performance.

The period of 1970s-90s was associated with other policy changes in the banking industry which reduced the obstacles for existing banks to open new branches within and between U.S. states. As shown in Table 1, Panel B, U.S. states deregulated restrictions on intrastate banking during the mid-1970s and 1980s. Further, the Riegle-Neal Interstate Banking and Branching Efficiency Act (IBBEA) enacted a nation-wide deregulation of the banking sector. As of 1995, the IBBEA removed any remaining federal restriction on interstate *banking*. Moreover, the IBBEA permitted national or state banks to engage in the interstate *branching*, although it also allowed states considerable leeway in deciding the rules governing entry by out-of-state branches (Rice and Strahan 2010; Johnson and Rice 2008).

One concern with our identification strategy is that deregulations may have been correlated with pre-existing trends in financial and economic development in the legislating states. If that is the case, our empirical approach would simply reflect pre-deregulation trends rather than an increase in innovation due to the exogenous changes in credit markets. We rule out this concern in three ways. First, we draw on the political economy of deregulations. Interstate banking deregulations were partly driven by the savings and loan crisis in the early 1980s, after which federal legislators enacted the Garn-St Germain Act. One provision of this act authorized federal banking agencies to arrange interstate acquisitions for failed banks with total assets of over $500 million, even when such acquisitions were not in accordance with state law. These changes paved the way for bilateral and regional negotiations between states to allow interstate banking and thus the creation of larger and more diversified banks that were less susceptible to failure (Kerr and Nanda 2009). Second, we empirically show that deregulations did not have a significant effect on innovations in the years prior to the actual deregulation passages (see Table 5, Panel A) and thus our estimates do not merely reflect pre-deregulation trends. Third, in Figure 2 we show that the association between pre-deregulation growth in patenting, computed as state averages over the years prior to deregulations, and the timing of deregulations are not significantly correlated. The *t* statistics of the slope obtained from an OLS regression is 1.01.

An additional concern is that banking deregulations might have no effect of innovative activity since a debt contract might be ill-suited to finance such activities as innovation that have uncertain returns (Atanassov et al. 2007; Stiglitz 1985). Yet, public firms might resort to private debt to fund innovation when they incur costs to raise capital in public markets. Indeed, funding innovation with public capital might provide sensitive information to the competitors (Bhattacharya and Ritter 1983; Maksimovic and Pichler 2001), or it can be costly to manager because of low tolerance for failure in the public markets (Ferreira et al. 2011).

With the data at hand we are not able to establish that corporate innovations rose because the affected banks directly increased lending to innovating firms. Alternatively, banks could have fostered financial development in the deregulating state and this had an indirect effect on the availability of finance for firms in that state. For instance, individuals could have increased their borrowing capacity and due to home bias channeled their portfolio investments to the nearby firms' equity offerings. Finally, although we study publicly listed firms, bank financing for innovations might have increased for private firms and this could have had innovation spillovers for larger, publicly listed firms. Either way, due to data limitations we abstract from how increased funds are actually channeled from banks to firms and rather focus on establishing that banking development has a positive effect on the corporate innovative activity.

3. Data and summary statistics

We start by collecting firm-level data from the Compustat dataset, which contains a rich set of financial characteristics for U.S. publicly traded firms. We measure firm innovation activities by successful patent applications, which represent a widely-used approach to quantify a firm's innovative performance (Griliches 1990).

Figure 1 shows the state-level non-parametric relationship between U.S. patenting activity, from the U.S. Patent and Trademark Office (USPTO), and loan supply data from the FDIC. Despite the presence of a strong positive correlation, it is challenging, to attribute to this evidence a causal interpretation. We establish the direction of causality by adopting the passage of interstate banking deregulations as our source of identification.

We focus the analysis on the period 1976-1995, which covers all years when interstate banking rules were deregulated but also include a few years before the passage of the first interstate banking deregulation, in 1978. We do not extend our sample after 1995 since in this year the interstate *banking* provisions of the IBBEA went into effect. Also, ending our sample in the mid-1990s ensures that our findings are not contaminated by the passage of state laws surrounding the IBBEA implementation which affected the evolution of interstate *branching* deregulations from 1994 to 2005 (Johnson and Rice 2008; Rice and Strahan 2010).[4] In a number of robustness checks, we show that our results do not change if we end our sample in 1994 (i.e. one year before the IBBEA enacted the nation-wide interstate *banking* deregulation) or in 1997 (i.e. the year when the IBBEA enacted the nation-wide interstate *branching* deregulation).

Following the literature on U.S. banking deregulations, we exclude observations in Delaware and South Dakota because these states were subject to special tax incentives. We also exclude firms with negative or zero book value of assets and sales, and firms headquartered outside the U.S.[5] As documented in Scherer (1983) and more recently in Balasubramanian and Sivadasan (2011), the bulk of patenting activity occurs within the manufacturing sector. Thus, following e.g. Hall and Ziedonis (2001) and Hall et al. (2005), we only consider firms in the industries with SIC codes up to 4000 (mostly manufacturing firms). Therefore, we exclude industries such as financial services or utilities, which typically operate under specific regulations, as well as the software industry.[6] Since the latter is primarily dependent upon non-debt sources of financing, such as equity issuances and venture capital, we believe that our identification strategy is less applicable for identifying the finance-innovation nexus in the software industry.

[4] Another advantage of ending the sample in 1995 is that our analysis is not contaminated by the dramatic increase in cash flow and equity financing of R&D activities experienced by young U.S. firms during the second half of the 1990s (Brown et al. 2009).

[5] A concern arises from the fact that Compustat only reports the last state of operations, and we may be unable to observe changes of headquarter from a state to another that are potentially endogenous to the deregulation passages. However, using data on headquarter relocations from the Compact Disclosure database, Pirinsky and Wang (2006) argue that most of the headquarter changes are driven by mergers and acquisitions. After excluding these and other major restructuring events, they found 118 relocations from a sample of more than 4000 firms in the period 1992-1997. Our results are robust to excluding firm-year observations with asset or sales growth exceeding 100%, which are typically associated with mergers, restructuring and other major corporate events (Almeida et al. 2004).

[6] In unreported analyses, we included in the analysis wholesale and retail trade sectors (SIC 5000-5999).

We match Compustat firms with the patent dataset assembled at the National Bureau of Economic Research (NBER), which contains information on the patents awarded by the USPTO and all citations made to these patents (Hall et al. 2001; Bessen 2009).

Table 2 reports summary statistics for the sample we use in the empirical analysis, obtained after dropping observations with missing values in the explanatory variables described in the next section. As documented in previous works on the Compustat-NBER patent dataset, citation statistics are very skewed. In our sample, the average number of patents is approximately 10 but the median is 1. The detailed construction of all control variables is described in the Appendix.

4. Methodology

We use a difference-in-differences model to explore the causal relationship between firm innovation and interstate banking deregulations. The important advantage of using a difference-in-differences model is that we can control for omitted variables and absorb nation-wide shocks or common trends that might affect the outcome of interest.

We expect a positive effect on innovation to arise from interstate banking deregulations because, in addition to an overall increase in credit supply, banks improved their scope for geographic diversification of credit risk, thereby becoming better able to lend to riskier companies. We use a binary variable $Interstate\ deregulations_{jt}$ which is equal to one if a firm i is headquartered in a state j which has passed an interstate banking deregulation at time t, and zero otherwise.[7] Hence, $Interstate\ deregulations_{jt}$ captures the effect of interstate banking deregulations on firm patenting by comparing outcomes before and after each deregulation year *vis-à-vis* deregulations passed later. Our first approach consists in using the logarithm of patent counts as dependent variable and estimating OLS regressions.

[7] As shown by Bharath et al. (2007), public firms have a strong propensity to borrow from local lenders, so we assume that firms should be primarily affected by the banking deregulations in the state of their headquarters.

However, this approach does not appropriately deal with firms that have zero patents. To avoid omitting these observations, we employ count data models, which are widely used in the econometric analysis of patents. Following Hausman et al. (1984), we hypothesize that the expected number of patents of a firm *i* applied for in year *t* is an exponential function of the intestate deregulation treatment, *Interstate deregulations*$_{jt}$, and firm- and industry-specific characteristics. More specifically, we estimate conditional-mean Poisson models:

$$E\left[Y_{ijt}\middle|Interstate\ deregulations_{jt}, X_{it-1}\right] =$$
$$ex\,p\left(\alpha + \beta Interstate\ deregulations_{jt} + + \delta X_{ijt-1} + \eta_i + \tau_t\right) \quad (1)$$

The model is estimated by the method of Quasi-Maximum-Likelihood (QMLE), which provides consistent estimates as long as the conditional mean is correctly specified even if the true underlying distribution is not Poisson (Wooldridge 1999). Since our deregulation treatment is defined at the state level, we cluster standard errors at the state of location.

Following the literature on the production function of patents (see e.g. Galasso and Simcoe 2011; Aghion et al. 2009), our baseline specification includes a vector X_{ijt-1} of time-varying controls, such as firm sales[8] and capital-labor ratio, all lagged by one year to reduce simultaneity concerns. Furthermore, we control for firm fixed effects denoted by η_i.

In additional analyses, we include other one-year lagged controls such as firm age and asset tangibility, to control for existing dependence and access to bank credit, return on assets (ROA), to control for profit positions, and cash holdings, to control for differences in liquidity. In addition, we construct the Herfindahl-Hirschman Index (HHI) to control for the impact of industry concentration on innovation.[9] The HHI is based on the distribution of revenues of the firms in a particular three-digit SIC industry. A higher HHI implies a higher concentration. We correct for potential misclassifications due to the presence of a single

[8] In unreported analyses, we control for firm size in alternative ways, such as by including the square of firm sales, or replacing firm sales with the logarithm of total assets.
[9] In unreported analyses, we check that our results hold even after including the squared HHI to capture potential non-linear effects of competition on innovation (Aghion et al. 2005).

firm in a given industry by dropping 2.5% of the firm-year observations at the right tail of the HHI distribution (Giroud and Mueller 2010).

Depending on the specification, we also control for the stock of R&D.[10] As stressed by Aghion et al. (2009), not controlling for the R&D stock implies that the coefficient of the variable of interest on the right-hand side will reflect both the increase in R&D expenses and the productivity of R&D. By contrast, when the R&D stock is included in the specification, the effect of the variables of interest can be interpreted as an effect on the innovative productivity of firms.

Given that the U.S. patenting activity increased substantially starting from the mid-1980s (see e.g. Hall 2004), it is important to control for aggregate trends. First, we include a full set of year dummies, denoted by τ_t in equation (1). Second, we control for industry linear trends by means of the annual three-digit SIC industry averages of the dependent variables, computed excluding the firm in question. Furthermore, in unreported analyses we check that our results are robust to including regional linear trends, and polynomial terms of the industry and regional linear trends.

Finally, we control for confounding policy changes. For instance, as discussed in Section 2, in the same period U.S. states deregulated interstate and intrastate branching activities. We conduct a number of robustness check showing that our results are not confounded by these policy changes.

5. Innovation activity

We further present our empirical results. First, we show the effect of banking deregulation on the number of patents successfully filed by the firms. Second, we analyze the quality of innovation by exploiting information on the citations that each patent received from

[10] The R&D stock is computed following the conventional 15% depreciation rate used in the related literature (see e.g. Hall et al. 2005). Also, we use linear interpolations to replace missing values of R&D; however, our results are robust to leaving those observations missing or treating them as zeros.

subsequent patent applications. We proceed by estimating the dynamic effects of banking deregulations, and discussing a number of robustness checks.

5.1 Innovation outputs

Table 3, Panel A, shows OLS estimates with the logarithm of patents as the dependent variable. As reported in Column (1), interstate banking deregulations had a positive effect on the innovation outputs. In particular, allowing out-of-state banks to enter the state increased innovation of a firm located in the deregulating state by 17%.

While in Column (1) we only control for industry and year fixed effects, in Column (2) we also control for the logarithm of sales and capital-to-labor ratio, while in Column (3) we further control for the stock of R&D. As expected, the stock of R&D has a positive and significant effect on patenting; however, the deregulation coefficient remains significant at 1%. In Column (4), we confirm our findings by including a host of industry- and firm-level factors that may potentially affect innovation, such as HHI, firm age, ROA, tangibility and cash holdings (coefficients are unreported to save space). As shown, the deregulation coefficient remains both statistically and economically relevant, indicating a 17% increase in patenting.

In Columns (5)-(8), we adopt a more restrictive specification where we include firm fixed effects. As expected, restricting the identification to within-firm variations leads to a sensibly smaller magnitude of the deregulation coefficient, but its statistical significance is confirmed at 1% level. In the most restrictive specification (Column 8), the economic magnitude of the effect is a 12% increase in patenting for an average firm that experienced interstate banking deregulations.

In Table 3, Panel B, we provide estimates from fixed-effect Poisson QMLE regressions, which take into account that patent counts are skewed to zero.[11] Similar to our OLS results, the most restrictive specification (Column 4) indicates a 12% increase in patenting following interstate banking deregulations.

5.3 Dynamic effects

The real consequences of interstate banking deregulations on credit markets caused by the *actual* entry of banks in the new states may manifest over several years following the passage of deregulation legislations. Also, filing a patent is the outcome of an innovation process that might take some years. We thus validate our identification strategy by looking at how the patenting activity evolved dynamically after the deregulations were enacted.

We test for dynamic effects by drawing on specifications similar to Kerr and Nanda (2009). First, we construct a dynamic difference-in-differences model employing a set of dummies that measure the distance in years from each deregulation passage, using as reference group the period three years or earlier before deregulations. Results, reported in Table 4, Panel A, show that the coefficient prior to the interstate deregulation is small and statistically insignificant, thus confirming that our results are not driven by diverging pre-deregulation trends. By contrast, the post-deregulation coefficients are all positive and significant at conventional levels. Importantly, they become larger as we move forward from the reform year, with the largest effect corresponding to six and seven years after interstate banking deregulations.

Second, we allow the effect of deregulations on innovation to linearly grow over time using a variable equal to zero up to the deregulation year and then equal to the number of

[11] An alternative approach could be to use transformations of the dependent variable to avoid losing observations with zero patents (as happens in Table 3, Panel A). For example, in unreported analyses we have estimated OLS regressions using the logarithm of (1+patents) as dependent variable. However, given that these transformations are arbitrary and often not robust to alternative methods, we prefer to adopt count data models.

years since a deregulation was passed, capping the treatment effect at 8 years. Results, reported in Table 4, Panel B, confirm that interstate deregulations had a positive growing impact on firms' patenting activity.

5.4 Robustness checks

In Table 5, we present a number of robustness checks. We start by excluding firms headquartered in California and Massachusetts. During the period considered, these states, which account for 24% of observations in our sample, experienced a dramatic increase in both innovation activity and alternative financing sources, such as venture capital. Column (1) shows that excluding these states does not materially affect our main result.

As Kerr and Nanda (2009) show, interstate banking deregulations fostered the creation and closure of firms. Hence, one concern is that our results can be driven primarily by nascent innovative firms. We assess this issue, first, by restricting our analysis to firms that were present in the sample since the first year of our sample, i.e. 1976 (Column 2). Second, we also remove exit effects by restricting the analysis to firms that remain in the sample from 1976 to the last year, i.e. 1995. As shown in Column (3), adding this further restriction does not significantly alter our estimates, despite the large drop in sample size.

In Column (4), we provide a more restrictive specification which controls for state-level macroeconomic variables that are potentially correlated with both credit availability and innovation activities. Specifically, in addition to the usual controls used in Table 4, Column (4), we control for one-year lagged GDP growth and logarithm of population, obtained from the U.S. Bureau of Economic Analysis (BEA). Our results are only marginally affected by the inclusion of these controls.

Next, we control for geographic trends. As shown in Column (5), our findings are unchanged if we augment our specification with regional trends, computed as year averages of the dependent variables by region excluding the firm in question.[12]

Furthermore, we show that our results are not affected by the time period considered. As discussed in Section 2, the interstate banking and branching provisions of the IBBEA enacted a nation-wide deregulation of banking activities. In our main sample we consider as the last year 1995, i.e. the year when the interstate banking provisions of the IBBEA were implemented. Yet, in Columns (6) and (7) we show that our results do not change if we end our sample in 1994 (i.e. one year before the nation-wide interstate banking deregulation enacted by the IBBEA) or if we extend the sample up to 1997 (i.e. the year when the nation-wide interstate branching deregulation enacted by the IBBEA). In unreported regressions, we also check that our results hold if we start the analysis in 1978, i.e. the first interstate banking deregulation year.

Finally, we address the concern that other policies potentially affecting innovation were adopted around the same period of the interstate banking deregulations. As discussed in Table 1, Panel B, U.S. states also deregulated intrastate branching activities during the period considered in our analysis. A concern for our analysis is that U.S. states may have deregulated intrastate branching and interstate banking at the same time, or within a few years, and this overlap of different deregulations may bias our identification. In Column (8), we show that our results are unchanged if we augment our specification with an indicator controlling for the effect of intrastate branching deregulations on patents.[13] As discussed in Section 2, the IBEEA granted the right to erect barriers to out-of-state

[12] Regions are defined according to the four-grouping classification provided by the U.S. Census: west, midwest, northeast and south (http://www.census.gov/geo/www/us_regdiv.pdf). Re-examining the findings in Black and Strahan (2002), Wall (2004) shows that the effect of deregulations on entrepreneurship was positive in some U.S. regions but significantly negative in others. If we estimate region-specific deregulation effects, we find that the interstate deregulation coefficients are all positive, though their statistical and economic significance is not homogeneous across U.S. regions.

[13] This effect (untabulated) is negative and borderline significant at 5%.

branching, which affected the evolution of interstate branching deregulations up to 2005 (Johnson and Rice 2008; Rice and Strahan 2010). In Column (9), we extend our sample to 2006 and include the index developed by Rice and Strahan (2010) as further control. As shown, our coefficient of interest remains economically relevant and its statistical significance is confirmed at the 10% level.

In the late 1980s, thirty U.S. states passed a set of business combination (BC) laws that reduced the threat of hostile takeovers, thus weakening the governance role of the market for corporate control (Giroud and Mueller 2010; Bertrand and Mullainathan 2003). These laws might have influenced our results if corporate governance affected the managerial incentives to innovate[14], and that effect would not be captured by our specification since BC laws affected firms at their state of incorporation. To mitigate this concern, we include in the specification a dummy equal to one if firms were incorporated in the states that passed a BC law, from the year of the passage onwards, and zero otherwise (Column 10). Furthermore, in untabulated regressions we interact BC laws with interstate banking deregulations dummy, to allow for heterogeneous effects of deregulation on innovation depending on whether the firm was subject to BC laws. Our estimates indicate that the positive effect of banking deregulations on firm innovation is not affected by the changes in corporate governance induced by BC laws.

6. Innovation quality and risk

So far our results indicate that firms subject to interstate banking deregulations patented more innovations. In this section, we test whether not only the number of patents has increased but their average quality has risen as well, i.e. whether the effect did not purely

[14] The effect of corporate governance on innovation is, in fact, ambiguous. Some empirical studies indicate that worse corporate governance reduces the incentives to innovate (Atanassov 2009). Chemmanur and Tian (2011) argue that some degree of managerial entrenchment isolates CEOs from short-term pressures, thus inducing them to focus on long-term value creation and innovate more. Sapra et al. (2011) show that the effect of corporate governance on innovation follows a U-shaped relationship.

come from larger credit supply and thus lower rationing of projects being financed. We posit that the average increase in the quality of innovations stems from a rise in the risk of innovative projects being financed and greater willingness of banks to take credit risk. Furthermore, we test whether the increase in patents corresponded to a more ambitious and more risky innovation policy. Finally, we explore the bank side by testing whether the deregulation effect was stronger in the states where entering banks could better diversify their credit risk.

6.1 Innovation quality

Existing research has demonstrated that patents differ greatly in "value" and that simple patent counts, as the ones we adopt in the previous sections, do not capture the relative importance of the underlying inventions (Harhoff et al. 1999). In this section, we measure innovation quality by weighing each patent using the number of future citations received from subsequent patents (Trajtenberg 1990). Forward citations reflect the economic and technological "importance" as perceived by the inventors themselves (Jaffe et al. 2000) and knowledgeable peers in the technology field (Albert et al. 1991). Because forward citations suffer from truncation problems, we weight patent counts by truncation-adjusted cite counts from the NBER dataset (see e.g. Hall et al. 2001; Hall et al. 2005).[15]

Results reported in Table 4 indicate that interstate banking deregulations led to a significant and economically relevant increase in the quality of patenting. Results are consistent irrespective of whether no control variables apart for firm- and time-fixed effects are used (Column 1); whether the logarithm of sales and capital-to-labor ratio are included as the controls (Column 2); whether, additionally, R&D stock is included as the control (Column 3); and whether other confounding industry- and firm-level variables, such as

[15] The problem arises from the fact that "citations to a given patent typically keep coming over long periods of time, but we only observe them until the last date of the available data" (Hall et al. 2005). Besides the use of truncation-adjusted citation counts, the problem is mitigated by the inclusion of year fixed effect. In fact, our results are robust to the adoption of unadjusted citation counts.

HHI, firm age, ROA, tangibility and cash holdings are also controlled for (Column 4). The deregulation coefficient remains both statistically and economically relevant, indicating an average 10% increase in cite-weighted patent counts.

6.2 Technological nature and risk of innovation

In this section, we investigate the technological nature and riskiness of firms' patenting activity. First, we combine citations with information on patents' technological fields. Second, we check if there is a simultaneous increase in both high-quality and low-quality patents. Finally, we analyze the volatility of successful patenting.

Technological fields, defined by the USPTO, consist of about 400 main (3-digit) patent classes. We use the *generality* and *originality* indexes, developed by Trajtenberg et al. (1997) and computed by Hall et al. (2001), to capture the fundamental nature of the research being patented. The generality index is equal to $1 - \sum_{j}^{n_j} s_{ij}^2$, where s_{ij}^2 denotes the percentage of citations received by a patent *i* that belong to the patent technology class *j* out of n_i patent classes. The index will take high values (high generality) if a patent receives citations from subsequent patents that belong to many different technological fields. The originality index is constructed in a similar way, but its computation relies on the citations made rather than citations received; it will take a high value if a patent cites other patents that belong to many different fields (high originality).

We use these two indices as the dependent variable in separate specifications similar to the ones used in Table 4. As reported in Table 7, interstate banking deregulations had a positive and significant effect on the generality of patents: firms subject to the interstate treatment exhibited a higher propensity to patent within broader technological fields (Columns 1 and 2). Moreover, the firms increased the originality of patents (Columns 3 and 4). Taken together with our previous finding on citations, these results suggest that a wider access to external finance led to a more ambitious innovation policy.

At the same time, a more ambitious innovation policy may entail more potential failures. Analyzing the distribution of cite-weighted patent counts, our additional analyses show that firms' successful patenting indeed became riskier. First, we study whether there was a simultaneous increase in patents that received many citations as well as few citations in the future. We estimate quantile regressions at different percentiles in the distribution of the logarithm of cite-weighted patent counts. In line with our notion of increased risk, our results, reported in Table 8, Panel A, show that the effect is present both at low deciles (e.g. 30% and 40%) and high deciles (e.g. 80%).

Second, we analyze the volatility of successful patenting. Specifically, we adopt as the dependent variable the standard deviation of the logarithm of cite-weighted patent counts computed in the pre- and post-interstate deregulation periods, restricting the analysis to firms that are present at least two years in each period. We estimate a regression including the interstate deregulation dummy, the usual controls averaged over the pre- and post-deregulation periods, and the firm fixed effects. Results reported in Table 8, Panel B, indicate that interstate banking deregulations led to an increase in the volatility of successful patenting.

7. Channels

7.1 Banks' geographic diversification

One of the channels that can explain higher and riskier corporate patenting after the entrance of new banks is that out-of-state banks were better able to finance riskier projects as they were less exposed to the background risks of the state's economy. At the same time, credit in this state provides out-of-state banks an opportunity to diversify their loan portfolio, for instance, due to a different industry composition of the state. We use three empirical tests that provide empirical support to this argument.

In our first test, we separate the states according to how their economic activity comoves with the rest of the U.S. economy. Here we expect that states that are least correlated with the activity of other states would provide highest diversification benefits for entering banks and thus would experience highest increase in the number of patents. In particular, we extract a coincident index that summarizes state-level economic indicators from the Federal Reserve Bank of Philadelphia. The coincident index combines data on nonfarm payroll employment, average hours worked in manufacturing, unemployment rate, and wage and salary disbursements deflated by the consumer price index (U.S. city average). The trend for each state's index is set to the trend of its gross domestic product (GDP) so that long-term growth in the state's index matches long-term growth in its GDP (Crone and Clayton-Matthews 2005). We estimate the correlation of a state's economy to the rest of the U.S. from the monthly values of the coincident indices of the states as well as the coincident index of the U.S. over 1979-84, before interstate banking deregulations started to come into effect. We call this variable *U.S./state correlation*. In Table 9, Columns (1) and (2), we show that the increase in patenting primarily rose in the states with the recent history of least covariation with the rest of the U.S.

Our second test relies on the location of banking institutions that enter a given state. We investigate whether the effect on innovation was highest in those states where new out-of-state banks were entering from the states least comoving with the state in question. In particular, for each pair of states we estimate the correlation of their monthly values of the coincident indices over 1979-84. We then calculate the weighted average of these comovement measures across all out-of-state banking institutions operating in the state, based on the location of their bank holding companies. As a weight for each institution, we use the assets it has in the state as a fraction of the total assets in the state held by out-of-state banking institutions. We estimate such a measure, decreasing in the actual diversification, for each state and year. We call this variable *Entering banks/state* *correlation*.[16] Our data on the banking institutions come from the Reports of Condition and Income (Call Reports) that provide information on the financial activities and ownership structures of each banking institution. All banking institutions regulated by the FDIC, the Federal Reserve, or the Office of the Comptroller of the Currency are required to file the Call Reports. Since this data is only available to us starting from 1986, we conduct the analysis on a subsample between 1986 and 1995. In Table 9, Columns (3) and (4), we report that the increase in patenting was evident mainly in the states that experienced the entry of the banks from the states with the least comoving economic indicators.

Our third test explores the differences in the diversification of banks in each state. If our argument is correct, those states that experienced the largest change in the average diversification of banks operating in their state, should have experienced a higher credit risk taking by banks and thus a subsequent larger increase in innovative activities. We use the data from the Call Reports to estimate the diversification of each bank. Our procedure follows three steps. First, for each bank we identify whether it is a subsidiary to some other financial institution, i.e. it is controlled by a bank holding company. Second, for each bank holding company we estimate the distribution of the total assets across states, based on where its subsidiaries are located. In particular, we estimate Herfindahl-Hirschman Index as our diversification measure (Goetz et al. 2011). Third, for each state we calculate the weighted average of these diversification measures across all banking institutions operating in the state. As a weight for each institution, we use the assets it has in the state as a fraction of the total banking assets in the state of all institutions. In summary, our measure, decreasing in the actual diversification, *Geographic diversification* is estimated as:

$$\sum_{i}^{N_i}\left(\frac{A_{ij}}{A_j}\sum_{j}^{K_i}\left(\frac{A_{ij}}{A_i}\right)^2\right)$$

[16] Correlation between *U.S./state correlation* and *Entering banks/state correlation* is 20%.

where A_{ij} denotes the total assets of all subsidiaries of bank i in state j; A_i denotes the total assets of all subsidiaries of bank i across all K_i states it operates; A_j denotes the total assets of all banks in state j.[17] A large increase in diversification might either mean that new out-of-state banks were well diversified but also that in-state banks better diversified across other states. In Table 9, Columns (5) and (6), we show that patenting mainly increased in the states with the largest diversification of banking activities.

7.2. Financial dependence and innovation inputs

If easier access to credit was a channel through which banking deregulations across states affected corporate innovation, we expect the effect to be more prevalent among firms that operate in industries requiring high external finance. We shed light on this notion by testing how our main finding vary depending on the industry-level reliance on external capital. To this end, we classify firms based on whether the industry in which they operate was above or below the across-industry median of external financial capital raised at the year of deregulation. Our measure is thus similar to Rajan and Zingales's (1998) proxy for an industry's financial constraints. We estimate it in two ways. First, we use balance sheet measures. In particular, we take the average across industry of the combined net change in equity and debt, normalized by the firm's book value of assets, before the year of deregulation. Second, we use the data in the SDC New Issues database and estimate financial dependence as the total proceeds from issuance of securities over the year divided by the book value of assets, before the year of deregulation.

The results, which are reported in Table 10, Columns (1)-(4), indicate that the positive effect of interstate banking deregulations on firm innovations is more pronounced for firms operating in industries that are highly dependent upon external finance.

[17] Correlation between *U.S./state correlation* and *Geographic diversification* is -10%, while correlation between *Entering banks/state correlation* and *Geographic diversification* is 0.

Turning the attention to firm-level characteristics, we further posit that our finding should be stronger for firms that were more dependent upon bank credit prior to deregulations. We investigate this issue by first considering firm age. Because old and well-established firms can typically access the public debt market or easily raise equity, they should not be influenced by changes in bank credit supply. By contrast, young firms, which are typically more financially constrained due to asymmetric information problems, are expected to respond to changes in bank credit. We focus on the subsample of firms that are present for less than 10 years in Compustat (Rajan and Zingales 1998; Cetorelli and Strahan 2006) at the time of the interstate banking deregulations, and firms that entered the sample after deregulations. As shown in Table 11, Columns (1) and (2), the effect of deregulations on firm patents was positive and statistically significant both for young and non-young firms. However, the economic magnitude of the effect is much larger among young firms: while young firms subject to interstate banking deregulations experienced a 20% increase in patents, the effect is 12% for non-young firms.

Second, we sort firms according to whether in 1985 they were assigned a long term bond rating by Standard&Poors.[18] By allowing firms to access public debt markets, a bond rating is related to lower credit constraints (Kashyap et al 1994; Almeida et al. 2004; Faulkender and Petersen 2006; Denis and Sibilkov 2010) and, consequently, lower responsiveness to changes in bank finance (Leary 2009). We construct two subsamples depending on whether a firm reports a bond rating and positive debt or whether a firm is not assigned to a rating or it has no debt. Table 11, Columns (3) and (4) show that the deregulation effect on innovation is marginally larger for firms experiencing a tougher to the public bond market.

Overall, these results suggest that the effect of interstate banking deregulations on corporate innovation was shaped by bank dependence: the effect was economically larger

[18] 1985 is the first year when the coverage of S&P ratings in Compustat started.

among firms that were younger and that had worse access to other segments of the credit market. This evidence is consistent with previous findings that bank credit is most relevant for less-established and informationally opaque firms (Hadlock and James 2002).

We further study the importance of financial constraints by constructing the KZ index and estimating our model separately for constrained (above-median KZ index) and unconstrained (below-median KZ index) firms. To compute the KZ-index, we follow Lamont et al. (2001) who use the original coefficient estimates of Kaplan and Zingales (1997). Results, reported in Columns (5) and (6), show that the effect of interstate banking deregulations was economically more relevant among constrained firms, though the statistical significance is present in both subsamples.

If firms innovated more due to an expansion of credit following deregulations, we should expect our results to be increasing in the intensity of post-deregulation innovation inputs. We explore this aspects by classifying firms depending on whether in the post-deregulation period they invested more than their industry peers in R&D expenditures. Results are reported in Columns (7) and (8). As expected, the effect of deregulations on patents is only present among firms that invested heavily in innovation inputs following banking deregulations.

8. Creative destruction and innovative activities

Recent papers have provided evidence suggesting that deregulations had a positive effect on Schumpeterian creative destruction triggering output reallocations. Our results so far have indicated that interstate banking deregulations had also significant effects on firms' innovative activities. The last step of our analysis is to provide an insight into the association between industry-level output dynamics and innovation after deregulations. We posit that the increased innovation stemming from deregulations fostered creative destruction and thus output reallocation towards the most efficient firms within an industry.

We estimate separate regressions as in our basic specification (1) for each SIC 2-digit industries and rank industries by how much the patenting activity in a specific industry was affected by banking deregulation (i.e. by the size of the deregulation coefficient). In particular, we can see that the link between banking deregulation and patenting was largest in Primary Metal Industries (SIC 33), Furniture and Fixtures (SIC 25) and Petroleum Refining and Related Industries (SIC 29). When we compare these estimates at the industry level to the future growth of the industry, we find a close relationship. Figure 3 plots the industry growth in value of shipments over 1995-2000 against the industry-level deregulation coefficients. A positive relationship, significant at 7%, implies that industries where banking deregulation had higher impact on patenting experienced higher growth after the deregulations than industries where banking deregulation had lower impact on patenting. For instance, five industries with the largest estimates grew on average by 4.9% annually over 1995-2000. On the other hand, five industries with the smallest estimates grew on average by 0.2%. Furthermore, we find no difference in growth rates between these groups of industries before the wave of deregulation, in 1980-1985, and also we find that the association fades away over the longer period.

9. Conclusion

While the relationship between economic prosperity and financial development has been widely debated, establishing the direction of causality remains a challenging task. We focus on firms' innovative performance as a driving force of technological progress and growth, and exploit the passage of interstate banking deregulations during the 1980s and 1990s in the U.S. to generate exogenous variations in credit supply. Interstate banking deregulations allowed banks to better diversify their loan portfolios, increased the availability and quality of credit, and were associated with the adoption of new screening and monitoring technologies.

Our results indicate that interstate deregulations played a beneficial role in spurring firms' innovation activities, as measured by patent-based metrics. Furthermore, we find that the effect was not imminent and was mainly driven by firms that were bank dependent and that operated in industries requiring high external capital. Finally, we provide evidence that the increase in firms' innovation activities is associated with a better ability of out-of-state banks to diversify credit risk thereby lending to riskier borrowers.

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Figure 1.
Relationship between innovation and credit supply

This graph shows the non-parametric (lowess smoothing) relationship between total net loan supply and patenting activity in the U.S. using state-year observations from the mid 1980s to mid 1990s. The line reports the local linear regression fit computed using a bandwidth of 0.8.

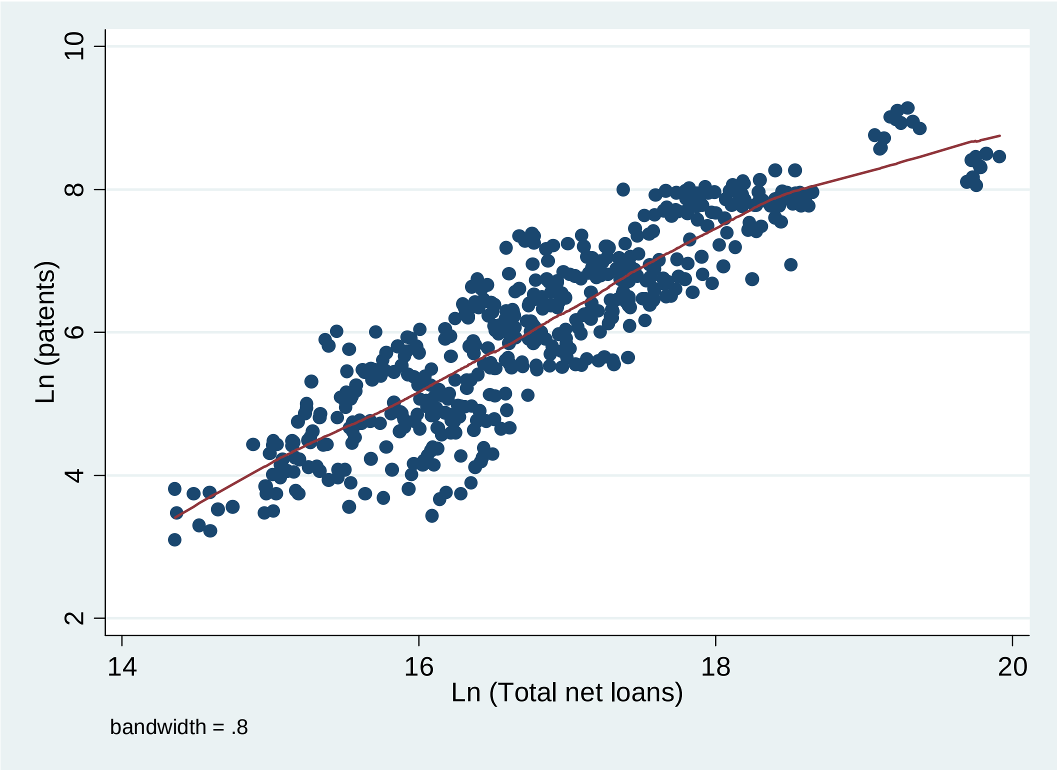

Figure 2.
Timing of interstate banking deregulations

This graph shows the relationship between the year of interstate banking deregulations and the patent growth in the pre-deregulation period computed as state averages. The line is the linear fit obtained from an OLS regression.

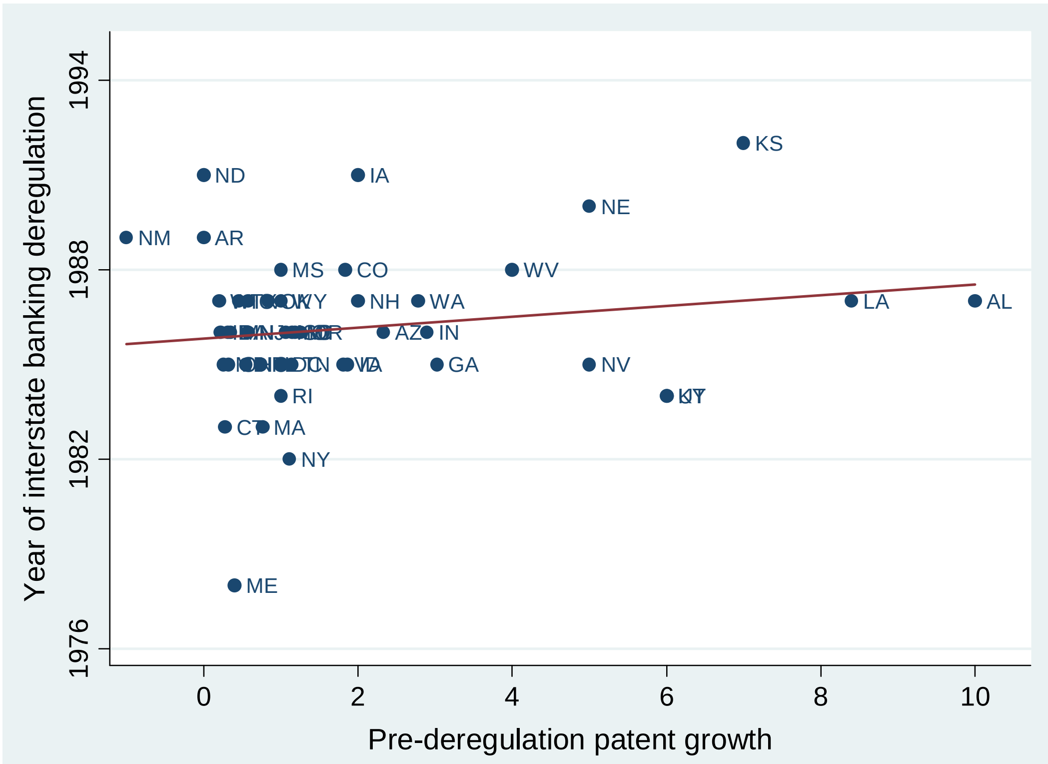

Figure 3.
Post-deregulation innovation and output by industry

This graph shows the relationship between the industry-level coefficients obtained from estimating specification (1) by 2-digit industry and the post-deregulation industry growth over the period 1995-2000. The line is the linear fit obtained from an OLS regression.

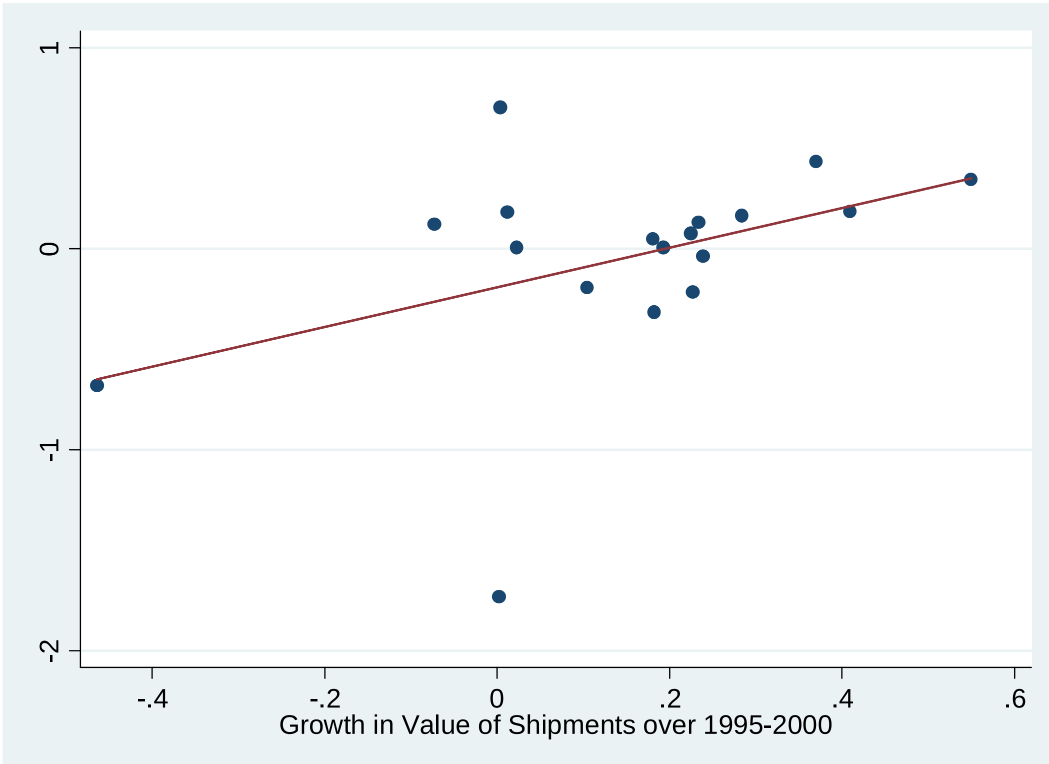

Table 1.
Major deregulations in the U.S. banking industry

Panel A. Interstate banking deregulations

Year	State
1978	Maine
1982	New York, Alaska
1983	Connecticut, Massachusetts
1984	Rhode Island, Utah, Kentucky
1985	North Carolina, Ohio, Virginia, District of Columbia, Nevada, Maryland, Idaho, Georgia, Tennessee, Florida
1986	Arizona, New Jersey, South Carolina, Pennsylvania, Oregon, Michigan, Illinois, Indiana, Missouri, Minnesota
1987	California, Alabama, Washington, New Hampshire, Texas, Oklahoma, Louisiana, Wyoming, Wisconsin
1988	Delaware, Vermont, South Dakota, Mississippi, West Virginia, Colorado
1989	New Mexico, Arkansas
1990	Nebraska
1991	North Dakota, Iowa
1992	Kansas
1993	Montana
After 1993	Hawaii

Panel B. Intrastate branching deregulations

Year	State
Before 1981	Maine, Alaska, Rhode Island, North Carolina, Virginia, District of Columbia, Nevada, Maryland, Idaho, Arizona, South Carolina, Delaware, California, Vermont, South Dakota, New York, New Jersey, Ohio, Connecticut
Between 1981 and 1985	Utah, Alabama, Pennsylvania, Georgia, Massachusetts, Tennessee, Oregon, Washington, Nebraska
Between 1986 and 1990	Mississippi, Hawaii, Michigan, New Hampshire, West Virginia, North Dakota, Kansas, Florida, Illinois, Texas, Oklahoma, Louisiana, Wyoming, Indiana, Kentucky, Missouri, Wisconsin, Montana
After 1990	Colorado, New Mexico, Minnesota, Arkansas, Iowa

Panel C. Interstate Branching and Banking Efficiency Act (IBBEA)

Provision	Implementation
Interstate banking	Took effect in September 1995; permitted the Board of Governors of the Federal Reserve System to approve interstate bank acquisitions, regardless of whether such acquisitions would have been permitted under "the law of any State." States were not permitted to opt out of the interstate banking rules.
Interstate branching	Took effect in June 1997; permitted national or state bank to engage in interstate branching. States could opt out of interstate branching or impose restrictions such as the minimum age of the target institution to be acquired or statewide deposit cap.

Table 2.
Summary statistics

This table illustrates summary statistics. Ln (Patents) is the logarithm of a firm's number of patents. Patent counts represent a firm's number of patents. Cite-weighted patent counts represent a firm's patents weighted by the number of future citations and adjusted for truncation. R&D to total investment is the ratio of R&D expenditures to the sum of R&D and capital expenditures. Ln (Sales) is the logarithm of a firm's sales. Ln (K/L) is the logarithm of capital to labor ratio. Ln (Age) is the logarithm of 1 plus the number of years a firm has been in Compustat. ROA is return on assets, measured as the ratio of earnings before interest and depreciation (EBITDA) divided by the book value of assets at the beginning of the year. See Appendix for a full description of each variable.

	Number of observations	Mean	Standard deviation	Median
Panel A. Innovation measures				
Ln (Patents)	11272	1.628	1.514	1.386
Patent counts	22400	10.418	40.681	1
Cite-weighted patent counts	22400	159.008	775.394	0
R&D to total investment	21688	0.427	0.266	0.400
Panel B. Main firm characteristics				
Ln (Sales)	22367	4.304	2.418	4.230
Ln (K/L)	22180	2.840	0.986	2.811
Ln (Age)	22400	2.525	0.785	2.565
ROA	22178	0.089	0.202	0.134

Table 3.
Innovation outcomes

This table reports regression results for number of patents. Panel A reports OLS regression results using Ln (Patents) as dependent variable while Panel B reports Poisson regression results using patent counts as the dependent variable. Columns (4) and (8) in Panel A and Column (4) in Panel B include an additional set of firm and industry controls. Specifically, they include one-year lagged Ln (Age), HHI, ROA, tangibility, cash holdings. Coefficients, unreported to save space, are available upon request. The construction of control variables is described in Appendix. Standard errors clustered by state of operation are reported in parentheses. *, ** and *** denote significance at 10%, 5% and 1% respectively.

Panel A. OLS estimates

Dependent variable: Ln (Patents)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Interstate deregulations	0.1755*	0.1596***	0.1852***	0.1758***	0.1352***	0.1288***	0.1219***	0.1193***
	(0.0881)	(0.0433)	(0.0458)	(0.0457)	(0.0373)	(0.0336)	(0.0367)	(0.0365)
Ln (Sales)		0.4494***	0.0849***	0.1369***		0.3437***	0.1786***	0.2112***
		(0.0218)	(0.0176)	(0.0250)		(0.0344)	(0.0468)	(0.0548)
Ln (K/L)		0.1406***	0.0438*	-0.0684**		0.0647	0.0224	-0.0087
		(0.0322)	(0.0231)	(0.0322)		(0.0444)	(0.0398)	(0.0459)
Ln (R&D stock)			0.4962***	0.4522***			0.3458***	0.3415***
			(0.0316)	(0.0357)			(0.0737)	(0.0792)
Industry fixed effects	Yes	Yes	Yes	Yes	No	No	No	No
Firm fixed effects	No	No	No	No	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry trends	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Additional controls	No	No	No	Yes	No	No	No	Yes
Number of obs.	11264	11264	11264	11264	11264	11264	11264	11264

Panel B. Poisson estimates

Dependent variable: Patent counts

	(1)	(2)	(3)	(4)
Interstate deregulations	0.1308**	0.1174**	0.1205***	0.1222***
	(0.0633)	(0.0471)	(0.0463)	(0.0409)
Ln (Sales)		0.7099***	0.4376***	0.4828***
		(0.0629)	(0.0769)	(0.0859)
Ln (K/L)		0.2546***	0.2366***	0.1871**
		(0.0635)	(0.0775)	(0.0801)
Ln (R&D stock)			0.4264***	0.3852***
			(0.1237)	(0.1244)
Firm fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Industry trends	Yes	Yes	Yes	Yes
Additional controls	No	No	No	Yes
Number of obs.	18011	18011	18011	18011

Table 4.
Dynamic effects

This table reports Poisson results using a dynamic specification. We use patent counts as dependent variable. In Panel A, the response to interstate and intrastate deregulations is modeled by leads and lags consolidated into two-year increments extending from two years before to eight years or more after the deregulations. Coefficients for leads are relative to the period three years or earlier before deregulations. Columns (3) and (4) include all controls as in Table 4, Column (8). Interstate and intrastate effects in Columns (1) and (2), as well as in Columns (3) and (4), are simultaneously estimated, although they are reported in separate columns to save space. In Panel B, years since interstate and intrastate deregulations are variables equal to the number of years after the deregulation passages, with a long-term effect at eight years, and equal to zero before the deregulations. Standard errors clustered by state of operation are reported in parentheses. *, ** and *** denote significance at 10%, 5% and 1% respectively.

Panel A. Time dummies

Dependent variable: Patent counts

	(1)	(2)
Years 1-2 before deregulation	0.0358	0.0424
	(0.0322)	(0.0263)
Years 2-3 after deregulation	0.1710***	0.1759***
	(0.0567)	(0.0430)
Years 4-5 after deregulation	0.1969**	0.2181***
	(0.0924)	(0.0816)
Years 6-7 after deregulation	0.2453**	0.2705***
	(0.1110)	(0.1033)
Years 8+ after deregulation	0.3904***	0.4129***
	(0.1335)	(0.1285)
Ln (Sales)	0.4121***	0.4355***
	(0.1123)	(0.1084)
Ln (K/L)	0.4942***	0.5321***
	(0.0790)	(0.0898)
Ln (R&D stock)	0.2445***	0.1988***
	(0.0762)	(0.0770)
Firm fixed effects	Yes	Yes
Year fixed effects	Yes	Yes
Industry trends	Yes	Yes
Additional controls	No	Yes
Number of obs.	18066	18066

Panel B. Linear treatment effects

Dependent variable: Patent counts

	(1)	(2)
Years since deregulation	0.0324*	0.0328*
	(0.0189)	(0.0179)
Ln (Sales)	0.4970***	0.5357***
	(0.0797)	(0.0905)
Ln (K/L)	0.2269***	0.1889**
	(0.0679)	(0.0745)
Ln (R&D stock)	0.3690***	0.3328***
	(0.1232)	(0.1232)
Firm fixed effects	Yes	Yes
Year fixed effects	Yes	Yes
Industry trends	Yes	Yes
Additional controls	No	Yes
Number of obs.	18066	18066

Table 5.
Robustness checks

This table reports Poisson regression results using patent counts as the dependent variable. Each column includes an additional set of firm and industry controls. Specifically, it includes one-year lagged Ln (Age), HHI, ROA, tangibility, cash holdings. Coefficients, unreported to save space, are available upon request. The construction of control variables is described in Appendix. Column (1) excludes firms headquartered in California and Massachusetts. Column (2) only includes firms present in the sample from 1976, i.e. the first year of our analysis. Column (3) further restricts the analysis to firms that remain in the sample until the last year of our sample, i.e. 1995. Column (4) includes state-level macroeconomic controls, such as the one-year lagged GDP growth and logarithm of population, both obtained from the U.S. Bureau of Economic Analysis. Column (5) includes regional trends, computed as annual average of the dependent variable by U.S. region (as defined by the U.S. Census), excluding the firm in question. Column (6) adopts a sample extended up to 1997, whereas Column (7) restricts the sample up to 1994. Column (8) includes as further control a dummy for intrastate branching deregulations, computed as the main interstate banking deregulation dummy. Column (9) extends our sample up to 2006 and further controls for the Rice and Strahan (RS) index of restrictions to interstate branching. Column (10) controls for anti-takeover legislations by including a dummy for the passage of BC laws. Standard errors clustered by state of operation are reported in parentheses. *, ** and *** denote significance at 10%, 5% and 1% respectively.

Dependent variable: Patent counts										
	Excluding California and Massachusetts	Excluding firm entry	Constant sample	Including state-level controls	Including regional trends	Time window up to 1997	Time window up to 1994	Controlling for intrastate deregulations	Controlling for RS index	Controlling for BC laws
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Interstate deregulations	0.1043***	0.0880**	0.1144***	0.1158***	0.1314***	0.1176**	0.1051***	0.1385***	0.1633*	0.1130***
	(0.0347)	(0.0398)	(0.0418)	(0.0399)	(0.0366)	(0.0529)	(0.0365)	(0.0395)	(0.0882)	(0.0433)
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry trends	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Additional controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Number of obs.	13891	11244	12333	18063	18063	20647	16969	18063	31658	16068

Table 6.
Innovation quality

This table reports Poisson regression results using cite-weighted and truncation-adjusted patent counts as the dependent variable. Column (4) includes an additional set of firm and industry controls. Specifically, it includes one-year lagged Ln (Age), HHI, ROA, tangibility, cash holdings. Coefficients, unreported to save space, are available upon request. The construction of control variables is described in Appendix. Standard errors clustered by state of operation are reported in parentheses. *, ** and *** denote significance at 10%, 5% and 1% respectively.

Dependent variable: Cite-weighted patent counts				
	(1)	(2)	(3)	(4)
Interstate deregulations	0.1334**	0.1010**	0.0988**	0.1006***
	(0.0651)	(0.0437)	(0.0391)	(0.0385)
Ln (Sales)		0.6895***	0.3185***	0.3733***
		(0.0571)	(0.0698)	(0.0915)
Ln (K/L)		0.2436***	0.2089***	0.1792**
		(0.0625)	(0.0776)	(0.0740)
Ln (R&D stock)			0.6005***	0.5807***
			(0.1055)	(0.1144)
Firm fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Industry trends	Yes	Yes	Yes	Yes
Additional controls	No	No	No	Yes
Number of obs.	17892	17892	17892	17892

Table 7.
Patenting and technological fields

This table reports Poisson results using as dependent variable the originality index (Columns 1-2) and generality index (Columns 3-4). The construction of these indexes and control variables is described in Appendix. Standard errors clustered by state of operation are reported in parentheses. *, ** and *** denote significance at 10%, 5% and 1% respectively.

Dependent variable:	Originality index		Generality index	
	(1)	(2)	(3)	(4)
Interstate deregulations	0.1041**	0.1043**	0.1358***	0.1368***
	(0.0516)	(0.0458)	(0.0505)	(0.0469)
Ln (Sales)	0.4441***	0.4856***	0.4639***	0.5118***
	(0.0686)	(0.0751)	(0.0718)	(0.0841)
Ln (K/L)	0.2415***	0.2100**	0.2251***	0.1581**
	(0.0852)	(0.0891)	(0.0647)	(0.0766)
Ln (R&D stock)	0.4034***	0.3630***	0.4654***	0.4234***
	(0.1190)	(0.1193)	(0.1317)	(0.1349)
Firm fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Industry trends	Yes	Yes	Yes	Yes
Additional controls	No	Yes	No	Yes
Number of obs.	17305	17305	16654	17305

Table 8.
Patenting risk

Panel A reports quantile regression results using Ln(Cite-weighted patent counts) as dependent variable. We include the full set of controls as in Table 4, Column (8). Coefficients, unreported to save space, are available upon request. Panel B reports OLS regressions using as dependent variable the standard deviation of Ln(Cite-weighted patent counts) in the pre- and post-interstate banking deregulation periods. Controls are constructed as average in the pre- and post-deregulation periods. *, ** and *** denote significance at 10%, 5% and 1% respectively.

Panel A. Distributional effects

Dependent variable: Ln (Cite-weighted patent counts)	
10th	-0.0966
	(0.1304)
20th	0.0148
	(0.0965)
30th	0.0931
	(0.0780)
40th	0.1588*
	(0.0927)
50th	0.1832**
	(0.0745)
60th	0.1752**
	(0.0697)
70th	0.1411**
	(0.0677)
80th	0.1417*
	(0.0752)
90th	0.0822
	(0.0749)

Panel B. Volatility effects

Dependent variable: σ (Ln (Cite-weighted patent counts))	
Interstate deregulations	0.1479**
	(0.0636)
Ln (Sales)	0.0077
	(0.0650)
Ln (K/L)	-0.0345
	(0.0642)
Ln (R&D stock)	-0.1109*
	(0.0652)
Year fixed effects	Yes
Industry trends	Yes
Number of obs.	866

Table 9.
Diversification benefits

This table reports Poisson regression results using number of patents as the dependent variable. U.S./state correlation refers to state economy's comovement with the rest of the U.S., measured as the correlation of state's coincident index to the U.S. coincident index. We estimate it from the monthly values of the indices over 1979-1984. The coincident index combines data on nonfarm payroll employment, average hours worked in manufacturing, the unemployment rate, and wage and salary disbursements deflated by the consumer price index (U.S. city average). Entering banks/state correlation refers to the weighted average of the comovements between the state and the states where the bank holding companies of its out-of-state banks are located. We estimate the pairwise correlations between all states from the monthly values of the coincident indices over 1979-1984. We then calculate the weighted average of these comovement measures across all out-of-state banking institutions operating in the state, based on the location of bank holding company. As a weight for each institution, we use the assets it has in the state as a fraction of the total assets in the state held by out-of-state banking institutions. Geographic diversification refers to the weighted average of diversification of all banking institutions operating in the state. As a weight for each institution, we use the assets it has in the state as a fraction of the total banking assets in the state. As a measure of each banking institutions' diversification, we estimate the Herfindahl-Hirschman Index based on the distribution of the assets of its subsidiaries across states. Due to data limitations, Entering banks/state correlation and Geographic diversification are only available for a subsample in the period 1986-1995, and that explains the smaller sample size in Columns (3)-(6). The construction of the control variables is described in Appendix. Standard errors clustered by state of operation are reported in parentheses. *, ** and *** denote significance at 10%, 5% and 1% respectively.

Dependent variable: Patent counts						
	High U.S./state correlation	Low U.S./state correlation	High entering banks/state correlation	Low entering banks/state correlation	Low geographic diversification	High geographic diversification
	(1)	(2)	(3)	(4)	(5)	(6)
Interstate deregulations	0.0257	0.1734***	0.1049	0.0973**	0.0320	0.1920***
	(0.0393)	(0.0818)	(0.1029)	(0.0425)	(0.0582)	(0.0509)
Ln (Sales)	0.3716***	0.4331***	0.2143***	0.5838***	0.5204***	0.3610***
	(0.1072)	(0.0975)	(0.0542)	(0.0934)	(0.1238)	(0.0709)
Ln (K/L)	0.2034*	0.2221**	0.3840***	0.1183***	0.1685	0.2812***
	(0.1069)	(0.0898)	(0.0892)	(0.1058)	(0.1289)	(0.0964)
Ln (R&D stock)	0.2238	0.5412***	0.5742***	0.3730**	0.4061*	0.4403***
	(0.1656)	(0.1672)	(0.1377)	(0.1785)	(0.2282)	(0.1164)
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Industry trends	Yes	Yes	Yes	Yes	Yes	Yes
Number of obs.	8213	9763	7956	10055	7532	10479

Table 10.
Industry-level financial dependence

This table reports regression results for number of patents. In Columns (1) and (2) financial dependence is estimated based on the balance sheet measures. In particular, we take the average across industry of the combined net change in equity and debt, normalized by the firm's book value of assets, before the year of deregulation. We then sort industries into high and low financial dependence based on the median financial dependence. In Columns (3) and (4), financial dependence is estimated based on the data in the SDC New Issues database and is equal to total proceeds from issuance of securities over the year divided by the book value of assets, before the year of deregulation. We then sort industries into high and low financial dependence based on the median financial dependence. All regressions include an additional set of firm and industry controls. Specifically, they include one-year lagged Ln (Age), HHI, ROA, tangibility, cash holdings. Coefficients, unreported to save space, are available upon request. The construction of control variables is described in Appendix. Standard errors clustered by state of operation are reported in parentheses. *, ** and *** denote significance at 10%, 5% and 1% respectively.

Dependent variable: Patent counts				
	High financial Dependence	Low financial dependence	High financial dependence	Low financial dependence
	(from balance sheets)		*(from new issuances)*	
	(1)	(2)	(3)	(4)
Interstate deregulations	0.1680**	0.0598	0.1502**	0.0816*
	(0.0664)	(0.0705)	(0.0590)	(0.0431)
Ln (Sales)	0.3349**	0.5936***	0.4870***	0.4827***
	(0.1532)	(0.0989)	(0.1886)	(0.1096)
Ln (K/L)	0.3484***	0.0113	0.3193**	0.0936
	(0.1070)	(0.1105)	(0.1384)	(0.0843)
Ln (R&D stock)	0.4550*	0.0547	0.2560	0.2612
	(0.2482)	(0.1163)	(0.2310)	(0.1741)
Firm fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Industry trends	Yes	Yes	Yes	Yes
Additional controls	Yes	Yes	Yes	Yes
Number of obs.	5593	6825	5860	6558

Table 11.
Bank dependence and innovation inputs

This table reports regression results for number of patents. Both Panel A and B reports Poisson regression results using patent counts as the dependent variable. In Panel A, Columns (1) and (2) report regression results separately for young firms and non-young firms. Young firms are defined as firms that were present for less than 10 years in the Compustat dataset at the time of the interstate banking deregulation or that entered the dataset after deregulation. Columns (3) and (4) report regression results separately for firms that had or not a bond rating in 1985. Columns (5) and (6) report regression results separately for firms with Kaplan-Zingales (KZ) score above or below the median value. Columns (7) and (8) report regression results separately for firms with Size-Age (SA) index above or below the median value. All regressions include an additional set of firm and industry controls. Specifically, they include one-year lagged Ln (Age), HHI, ROA, tangibility, cash holdings. Coefficients, unreported to save space, are available upon request. The construction of control variables is described in Appendix. Standard errors clustered by state of operation are reported in parentheses. *, ** and *** denote significance at 10%, 5% and 1% respectively.

Panel A. Financial constraints

Dependent variable: Patent counts

	Young firms	Non-young firms	Firms without bond rating	Firms with bond rating	Constrained (KZ index)	Unconstrained (KZ index)	Innovative firms	Other firms
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Interstate deregulations	0.2047*	0.1114***	0.1279*	0.1142**	0.1995*	0.0680*	0.1177***	-0.0195
	(0.1170)	(0.0418)	(0.0683)	(0.0541)	(0.1130)	(0.0366)	(0.0405)	(0.1603)
Ln (Sales)	0.1189**	0.5998***	0.0125	0.2372*	0.3352***	0.4823***	0.5553***	0.3990***
	(0.0555)	(0.1086)	(0.1035)	(0.1241)	(0.1170)	(0.0990)	(0.1032)	(0.0791)
Ln (K/L)	0.0053	0.2521***	0.3453***	0.7053***	0.1537	0.2123**	0.2136**	0.1021
	(0.0730)	(0.0954)	(0.0830)	(0.1174)	(0.1336)	(0.1031)	(0.0882)	(0.1291)
Ln (R&D stock)	0.6150***	0.2340	0.6305***	0.2994*	0.6407***	0.1599	0.2815**	0.4302***
	(0.1465)	(0.1540)	(0.0775)	(0.1787)	(0.0998)	(0.1341)	(0.1361)	(0.1601)
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry trends	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Additional controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Number of obs.	5796	8874	7932	4378	7824	7105	8910	6790

Appendix.
List of variables

Name	Description	Source
	Innovation variables	
Patent counts	Count of a firm's number of patents for the period 1976-1995	NBER
Ln (Patents)	Logarithm of a firm's number of patents for the period 1976-1995	NBER
Cite-weighted patent counts	Count a firm's number of patents for the period 1976-1995 weighed by future citations received and adjusted for truncation (as described in Hall et al. 2001; Hall et al. 2005)	NBER
Ln (Cite-weighted patent counts)	Count a firm's number of patents for the period 1976-1995 weighed by future citations received and adjusted for truncation (as described in Hall et al. 2001; Hall et al. 2005)	NBER
σ (Ln(Cite-weighted patent counts))	Standard deviation of the logarithm of a firm's count of number of patents for the period 1976-1995 weighed by future citations received and adjusted for truncation. Standard deviations are computed in the pre-and post-deregulation period, keeping in the computation firms that remain in each period at least two years	NBER
Originality index	Equal to $1-\sum_j^{n_j} s_{ij}^2$, where s_{ij}^2 denotes the percentage of citations made by a patent *i* that belong to the patent technology class *j* out of n_i patent classes. Technology classes are defined by the USPTO and consist of about 400 main patent classes (3-digit level). The index will take high values (high originality) if a patent cites other patents that belong to many different technological fields	NBER
Generality index	Equal to $1-\sum_j^{n_j} \mathrm{s}_{ij}^2$, where s_{ij}^2 denotes the percentage of citations received by a patent *i* that belong to the patent technology class *j* out of n_i patent classes. The index will take high values (high generality) if a patent receives citations from subsequent patents that belong to many different technological fields	NBER
Ln (R&D Stock)	Logarithm of (1 + cumulative R&D expenditures), computed assuming a 15% annual depreciation rate and using linear interpolation to replace missing values of R&D	Compustat
R&D to total investment	Ratio of R&D expenses to total investment, computed as the sum of CAPEX and R&D expenses	Compustat
	Firm and industry characteristics	
Ln (Sales)	Logarithm of a firm's sales	Compustat
Ln (K/L)	Logarithm of capital to labor ratio, where capital is represented by property, plants and equipment (PPE), and labor is the number of employees	Compustat
Ln (Age)	Logarithm of (1+age), where age is the number of years that the firm has been in Compustat	Compustat
ROA	EBITDA to total assets, dropping 1% of observations at each tail of the distribution to mitigate the effect of outliers	Compustat
Cash holdings	Cash and marketable securities to total assets	Compustat
Tangibility	1- (intangible assets to total assets)	Compustat
HHI	Herfindahl-Hirschman Index, computed as the sum of squared market shares of all firms, based on sales, in a given three-digit SIC industry in each year. We drop 2.5% of observations at the right tail of the distribution to mitigate potential misclassifications (Giroud and Mueller 2010)	Compustat
Young firms	Dummy equal to one if a firm was present for less than 10 years in Compustat at the time of the interstate banking deregulation, and zero otherwise	Compustat
Bond rating	Dummy equal to one if a firm report a S&P bond rating in 1985, and zero otherwise	Compustat
Industry trends	Average of the dependent variable across all firms in the same three-digit SIC industry of a given firm, where averages are computed excluding the firm in question	Compustat
	Banking deregulations variables	
Interstate deregulations	Dummy variables equal to one from the year of interstate banking deregulation year onwards, and zero for the years prior to deregulations	

Diversification variables

U.S./state correlation	State economy's comovement with the rest of the U.S., measured as the correlation of state's coincident index to the U.S. coincident index. We estimate it from the monthly values of the indices over 1979-1984. The coincident index combines data on nonfarm payroll employment, average hours worked in manufacturing, the unemployment rate, and wage and salary disbursements deflated by the consumer price index (U.S. city average)	Federal Reserve Bank of Philadelphia
Entering banks/state correlation	Weighted average of the comovement between the state and the states where the bank holding companies of its out-of-state banks are located. We estimate the pairwise correlations between all states from the monthly values of the coincident indices over 1979-1984. We then calculate the weighted average of these comovement measures across all out-of-state banking institutions operating in the state, based on the location of bank holding company. As a weight for each institution, we use the assets it has in the state as a fraction of the total assets in the state held by out-of-state banking institutions. Due to data limitations, this measure is constructed for the period 1986-1995	Federal Reserve Banks of Philadelphia and Chicago
Geographic diversification	Weighted average of diversification of all banking institutions operating in the state. As a weight for each institution, we use the assets it has in the state as a fraction of the total banking assets in the state. As a measure of each banking institutions' diversification, we estimate the Herfindahl-Hirschman Index based on the distribution of the assets of its subsidiaries across states. Due to data limitations, this measure is constructed for the period 1986-1995	Federal Reserve Bank of Chicago