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Abstract

This paper exploits the staggered timing of state-level banking deregulation in the United States during the 1980s to study the causal effect of banking integration on the volatility of non-financial corporations. We find that firm-level employment, production, sales, and cash flows are less volatile after interstate banking deregulation, particularly for firms that have limited access to external finance. This finding suggests that bank-dependent firms exploit wider access to finance after deregulation to smooth out idiosyncratic shocks. In fact, short-term credit becomes less pro-cyclical after out-of-state bank entry is permitted. Finally, lower volatility in real-side variables after deregulation translates into lower idiosyncratic risk in stock returns.

JEL Classification Codes: G21, G32

Keywords: Bank deregulation, firm volatility, external finance, idiosyncratic volatility

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A growing body of theoretical and empirical research indicates that the ability of firms to access banking finance fosters average growth.[1] For example, Rajan and Zingales (1998) find that industries that depend on external financing for investment grow faster in countries with larger banking systems. In addition to its positive effect on average growth, access to banking finance may have an important effect on growth volatility. The effect of financial markets—and banking finance in particular—on volatility is less well understood. This paper studies the relationship between banking integration and the volatility of the corporate sector using data on publicly-traded U.S. firms.

From a macroeconomic point of view, volatility is important because of the growth benefits conferred by stability (Ramey and Ramey (1995), Aghion et al. (2005)). From the point of individual corporations, volatility is important because stable firms face lower expected costs from financial distress (Smith and Stulz (1985)), financial policies are more effective for solving agency problems in stable firms (Stulz (1990)), and investors value firms with smooth cash flows at a premium (Rountree, Weston, and Allayannis (2008)).[2]

In theory, the direction of the effect of banking finance on firm volatility is ambiguous. On the one hand, wider access to external finance may increase the ability of firms to offset idiosyncratic shocks, thus reducing firm volatility. On the other hand, wider access to external finance may promote specialization and allow firms to pursue riskier and more profitable projects (Thesmar and Thoenig (2009)), thus increasing firm volatility. As a consequence of mutually offsetting forces, the net effect of access to banking finance on firm volatility is an empirical question.[3]

Recent papers in this area suggest that wider access to banking finance may lower volatility in the corporate sector. Extending the methodology of Rajan and Zingales (1998), Larrain (2006) finds that growth in industries that depend on external finance is less volatile in countries with more bank credit relative to GDP. Using a similar cross-country, cross-industry approach, Raddatz (2006) argues that the volatility-reducing effect of more bank intermediation results partly from the role of the financial system in

[1] See Levine (2005) for a survey.
[2] The existence of corporate hedging activity suggests that firms try to mitigate some sources of risk and volatility. See, for example, Froot, Scharfstein, and Stein (1993).
[3] In addition, the effect of wider access to banking finance on volatility may depend on the relative strength of credit supply and credit demand shocks (Morgan, Rime, and Strahan (2004)).

providing liquidity.[4] Cross-country studies results may also be consistent with a reverse causality interpretation, where a more stable corporate sector facilitates the expansion of bank credit.

This paper complements and refines the findings in Larrain (2006) and Raddatz (2006) by using a different identification strategy for isolating the effect of banking finance on corporate volatility. In particular, we use a differences-in-differences approach that exploits the staggered timing of interstate banking deregulation across U.S. states during the 1980s and early 1990s to estimate the effect of banking integration on firm-level volatility. After a state passed an interstate banking deregulation law, out-of-state banks were allowed to acquire banks incorporated in that state, effectively promoting the integration of banking markets. We interpret deregulation that facilitates banking integration, starting from a situation with highly segmented markets, as promoting wider access to banking finance and improving efficiency in intermediation, possibly reflecting that financial institutions become more geographically diversified (Jayaratne and Strahan (1996), Morgan, Rime, and Strahan (2004)). With this interpretation, the results of this paper may inform the debate about the impact of access to banking finance on firm-level volatility.

Our identification strategy is similar to Morgan, Rime, and Strahan's (2004), but we focus on firm volatility—as opposed to state volatility. Firm-level data help us in identifying banking deregulation as the driving force behind our results, because we are able to compare, within each state, the effect of deregulation on bank-dependent firms with the effect on non-bank-dependent firms. Previous research suggests that small firms and firms with limited access to corporate debt face significant asymmetries when accessing credit markets (Almeida, Campello, and Weisbach (2004), Gilchrist and Himmelberg (1995)). These types of firms tend to rely more heavily on banks for their external financing needs and are defined as *bank dependent* in our regression analysis. We expect bank-dependent firms to be more sensitive to changes in banking conditions within a state.

We find that firms located in states that deregulate interstate banking exhibit a

[4] In a related branch of the literature, Krozner, Laeven, and Klingebiel (2007) study the impact of banking crises across sectors depending on their degree of dependence on external finance.

reduction in the volatility of the growth rates of sales, production, cash flow, and employment after deregulation. In our main result, the reduction in volatility is likely to be associated with the changes in banking finance derived from deregulation, because it is concentrated on firms that are more likely to rely on banks for external financing. For example, firms without bond or commercial paper credit ratings or issues reduce their volatility after deregulation by more than firms with credit ratings or issues. Further tests suggest that firms exploit credit markets to smooth temporary cash-flow shocks, as short-term credit becomes more countercyclical after deregulation (as in Larrain (2006)).

In addition, by using firm-level data we are able to examine the implications of wider access to banking finance for the volatility of individual stock returns. We document that the idiosyncratic volatility of stock returns falls after deregulation, particularly for firms that are a priori more likely to rely on banking finance. The residual variance of a market model of excess returns falls after deregulation for firms located in states that open to interstate banking. This finding is robust to adding the size and value factors of Fama and French (1993) and the momentum effect of Jegadeesh and Titman (1993).

At first glance, our results might seem at odds both with the increase in volatility of firm sales growth documented by Comin and Philippon (2005) and Comin and Mulani (2006) and with the increase in the idiosyncratic volatility of individual stock returns documented by Campbell et al. (2001). While these papers describe aggregate trends over the second half of the twentieth century, we restrict our attention to the window of years around deregulation of interstate banking. Importantly, the increasing trend in volatility described by Comin and Philippon (2005) temporarily reverses during the 1980s—the years with the heaviest deregulation activity at the state level and, therefore, the main part of our sample. Similarly, the increasing trend in idiosyncratic volatility described by Campbell et al. (2001) does not apply to the 1980s.[5] In this paper, we abstract from aggregate trends in idiosyncratic volatility by controlling for year-fixed effects. We argue that part of the remaining variation in firm-level volatility is explained by interstate banking deregulation.

[5] In fact, excluding the market crash of 1987 in Figure 4 of Campbell et al. (2001), idiosyncratic volatility trended down during the 1980s.

The rest of the paper is organized as follows. Section I briefly reviews the interstate banking deregulation in the 1980s in the United States and summarizes some related studies. Section II presents our empirical methodology and the data. Section III shows the effect of banking deregulation on firm-level volatility and its differential impact on those firms that are more likely users of banking finance. This section also links volatility in operations and stock return volatility using multifactor models of stock returns at the firm level to isolate the idiosyncratic component of returns. Section IV concludes.

I. The Deregulation of Interstate Bank Entry

The Douglas Amendment to the Bank Holding Company Act of 1956 prohibited Bank Holding Companies (BHCs) from establishing or purchasing bank subsidiaries across state lines unless the state of the target bank authorized the transaction. These restrictions remained in place until Maine passed a law allowing out-of-state BHCs to purchase local banks if the "home" state of the BHC reciprocated. This did not happen until 1982 when Alaska and New York passed similar laws. The same year, as part of the Garn-St Germain Act, federal legislators amended the Bank Holding Company Act to allow failed banks to be acquired by any BHC, regardless of origin and state laws. This regulatory change, coupled with a series of bank and thrift failures during the eighties, triggered a wave of interstate agreements that effectively permitted banking at the national level. By 1994, 49 states and the District of Columbia had deregulated their banking markets allowing out-of-state entry.[6]

The episode of banking deregulation in the 1980s and early 1990s changed the terms in which nonfinancial firms were able to access banking finance. In particular, Jayaratne and Strahan (1998), Dick (2006), and Rice and Strahan (2009) find that loan prices and spreads decrease after banking deregulation.

A series of studies have analyzed the effect of this episode of banking deregulation on real economic activity. Strahan (2003) finds that interstate deregulation is associated with an increase in incorporations by state and a reduction in the link between state

[6] Interstate branching was permitted nationwide with the Riegle-Neal Interstate Banking and Branching Efficiency Act, which became effective in June 1997. Montana and Texas took advantage of a clause in the Act and opted out at an earlier date. The table in Appendix 1 lists deregulation dates by state.

growth and local bank performance.[7] In a study on income insurance, Demyanyk, Ostergaard, and Sorensen (2007) find that deregulation, measured as the combination of intrastate and interstate deregulation, decreases the correlation between personal income and state-specific shocks to output. Their result is stronger for proprietor income than for wage income. The authors explain this effect by the closer relationship between banks and small businesses. This result is connected to Morgan, Rime, and Strahan's (2004) finding that geographical bank integration reduces employment volatility within states. This change is due to a decline in the impact of bank capital shocks on state activity.

A limited number of studies have examined the connection between deregulation and firm-level dynamics. Correa (2007) finds that banking deregulation is associated with a decrease in financing constraints for small publicly-traded firms, explained by lower costs of financing after deregulation. From a theoretical perspective, Stebunovs (2006) uses a stochastic general equilibrium model to assess the consequences of banking deregulation on volatility. The author finds that deregulation increases firm entry by reducing local bank monopoly power, thus dampening firm- and aggregate-level fluctuations.

In this paper, we expand on these findings and analyze the empirical effect of interstate bank-entry deregulation on firm volatility. Moreover, we test if changes in the volatility of firm fundamentals are accompanied by changes in the volatility of individual stock returns.

II. Data and Methodology

A. Data

In the empirical analysis of this paper, we use balance sheet and income statement data extracted from the Compustat North-America database. The sample includes observations for firms classified in the manufacturing (SICs 2000 to 3999), wholesale (SICs 5000 to 5199), and retail trade (SIC 5200 to 5999) sectors. To focus on the episode of interstate banking deregulation of the 1980s and early 1990s, our sample includes data between 1976 and 1998.

[7] Huang (2007) exploits the variation of neighboring counties across state borders to understand the effects of this episode of banking deregulation.

Compustat's geographic information reports a firm's headquarters location only for the latest year available in the database. To determine whether a firm was affected by bank entry deregulation, we need to find the actual historical location of its headquarters. For this purpose, we collect data from Compact Disclosure between 1988 and 1998. This source contains extracts from SEC filings updated every month, including the firm's address. Using this information, we determine the state where the firm was headquartered during the deregulation period.[8] A firm is excluded from the sample if it changed its location between 1988 and 1998. In addition, we consider only firms with non-missing information two years before and two years after interstate bank entry deregulation in the state where headquarters are located.[9]

After applying these restrictions, the data consists of 39,624 observations for 2,272 firms in 45 states and the District of Columbia.[10] Table I shows the distribution of firms by state. Firms in New York and California account for one quarter of the sample; other important states in the Northeast (Connecticut, Maryland, Massachusetts, New Jersey, and Pennsylvania) account for about 20 percent of the sample; and large industrial states in the Midwest (Illinois, Michigan, and Ohio) represent 13 percent of the firms covered. Due to the sample selection criteria, an average firm stays in the sample for 17 years.[11]

[Table I]

Stock returns are collected from the Center for Research in Security Prices (CRSP) dataset. Additionally, the one-month Treasury Bill rate and the Fama-French and Jegadeesh-Titman factors are taken from Kenneth French's Data Library at Dartmouth.

[8] The earliest year for which we have location information comes after the first bank entry deregulation agreement was implemented (1982). Although, some firms might have changed their location between these dates, the number is likely to be small. Using information from Compact Disclosure, we determine that less than 10 percent of firms covered by this database changed the state where they were headquartered between 1988 and 2005.

[9] Due to this selection criterion, our results are explained by changes in the volatility of existing firms rather than changes in the volatility of entering and exiting firms. All results are robust to include firms that exit before deregulation and enter after deregulation.

[10] We exclude firms headquartered in Delaware and South Dakota due to their incorporation and consumer finance regulations, respectively. In addition, as Hawaii did not deregulate before 1994, we also exclude firms located in that state.

[11] The discontinuity at deregulation of our empirical design suggests that our results are not simply a reflection of firms becoming more stable as they age, and we conduct additional checks of our baseline results.

B. Firm Volatility

The effect of bank entry deregulation on firm volatility is ambiguous from a theoretical standpoint (Larrain (2006), Morgan, Rime, and Strahan (2004)). On the one hand, wider availability of bank credit may dampen the effect of idiosyncratic shocks on productive activities, making firms more stable. On the other hand, improved access to bank financing may increase volatility by allowing firms to undertake riskier and more profitable projects (e.g., adopting new technologies). Hence, we adopt an empirical approach to study the relationship between firm volatility and access to bank credit.

To estimate the effect of interstate bank entry deregulation on volatility we follow Morgan, Rime, and Strahan (2004) and use a two-stage procedure. In the first stage we calculate a time-varying measure of firm volatility. In the second stage we use this measure to determine whether interstate bank entry deregulation had an effect on the volatility of non-financial corporations.

Let i, j, k, and t index firm, state, industry, and year, respectively, and Y_{ijt} be a generic variable. We define the volatility of Y_{ijt} as the deviations in absolute value of Y_{ijt} from its predicted conditional mean. Formally, we estimate the following equation:

$$Y_{ijt} = \alpha_i + \mu_t + \phi Integration_{jt} + \eta X_{ijt-1} + \rho Z_{jt} + \kappa W_{kt} + \upsilon_{ijt} \quad (1)$$

where α_i are firm-fixed effects; μ_t are time effects; $Integration_{jt}$ is an indicator variable that equals one if state j permits out-of-state bank entry through acquisitions in year t; X_{ijt} are firm-specific controls; Z_{jt} represents the growth rate of state per capita income; and W_{kt} is the growth rate in real sales at the industry level (3-digit SIC). We consider three firm-level controls: the log of real sales (log(*Sales*)), tangible assets divided by total assets (*Tangible/Assets*), and earnings before taxes, depreciation, and amortization divided by total assets (*EBITDA/Assets*).[12] These measures are associated with firm-leverage (Rajan and Zingales (1995)), and reflect, respectively, the size of the firm, its collateral, and profitability.

[12] Following Almeida and Campello (2007), tangible assets are defined as: $Tangible = 0.715 \times Receivables + 0.547 \times Inventory + 0.535 \times Capital$, where *Receivables* is Compustat item #2, *Inventory* is item #3, and *Capital* is item #8. EBITDA is defined as Compustat item #13.

After estimating equation (1), the volatility for variable Y_{ijt}, $Vol(Y_{ijt})$, is defined as $\left|\hat{\upsilon}_{ijt}\right|$. Notice that this measure of volatility is the absolute deviation of the firm's indicators from the overall trends in the economy, from the state where its headquarters are located, and from the industrial sector that represents its main activity. In addition, we allow for changes in the firms' growth rate after interstate bank entry is permitted. The measure estimated from this empirical equation captures the idiosyncratic component of firms' volatility. Measuring volatility using absolute deviation (as opposed to squared deviations) implies that volatility and growth are conveniently expressed in the same units.

As part of their intermediation function, banks have a comparative advantage in providing liquidity to firms, particularly if the demand for funds is related to individual projects (Larrain (2006)). Therefore, to analyze the effect of banking deregulation on firm volatility, we will focus on its idiosyncratic component. In the second stage of our estimating procedure, we use the volatility measures defined in equation (1) and estimate the following equation:

$$Vol\left(Y_{ijt}\right) = \alpha_i + \mu_t + \delta Integration_{jt} + \beta X_{ijt-1} + \gamma Vol\left(Z_{jt}\right) + \varphi Vol\left(W_{kt}\right) + \varepsilon_{ijt} \qquad (2)$$

We add two regressors, in addition to the set included in equation (1), to control for fluctuation at the state and industry level. $Vol(Z_{jt})$ and $Vol(W_{kt})$ are volatility proxies for the states' per capita income growth and real sales growth at the industry level, respectively.

In our baseline regressions, Y_{ijt} is the growth rate of production, sales, cash flow, and employment, indicators of real corporate activity. These measures allow us to determine if output fluctuations changed after banking deregulation.[13] As in Larrain (2006), firm production is measured by inventory investments plus sales. In addition to this set of variables, we analyze the evolution of the firms' internal cash flow. Panel A in Table II shows summary statistics for the measures of volatility associated with these variables.[14]

[13] Nominal variables are deflated using the GDP deflator. Comin and Mulani (2005) test different price deflators, and do not find any significant differences in their volatility measures.

[14] Sales is Compustat item #12, employment is item #29, cash flow equals item #18 plus item #14, and inventories is item #3.

[Table II]

To test whether firm volatility decreases after interstate bank entry deregulation, we examine the sign of δ in equation (2). As documented by Comin and Mulani (2005), the volatility of publicly-traded firms increased over the last five decades. The inclusion of time-fixed effects captures this secular trend. Therefore, a negative sign on the *Integration* coefficient is interpreted as a decrease in the upward trend in volatility.

We control for variation at the industry (W_{kt}) and state level (Z_{jt}) to isolate the effect of deregulation on the idiosyncratic component of firm-specific volatility. The former is measured by the absolute value of deviations of the log change in sales at the 3-digit SIC level, *Vol*(*Industry Sales*).[15] The state-level component is proxied by the volatility of log changes in per capita income, *Vol*(*Per Capita Income*).[16] Finally, selecting firms with observations before and after deregulation and controlling for firm-fixed effects alleviates sample biases in which firms that enter after deregulation exhibit a systematically different volatility than firms that exit before deregulation.

C. Firm Volatility and Bank Dependence

The baseline empirical estimation described in the previous section takes advantage of the staggered deregulation dates across states to identify whether there was a change in the firms' volatility measures explained by lower restrictions on out-of-state bank entry. However, this strategy does not fully control for other potential shocks that might have had an effect on firms headquartered within the state at the time of banking deregulation. We take this problem into account by using another layer of differentiation. In particular, we test whether the effects of bank entry deregulation is stronger for bank-dependent firms, as opposed to firms with access to other sources of external financing, within each state.

We define two proxies for firms' *Bank Dependence*: one based on size and the other

[15] Formally, $Vol\left(Y_{kt}\right) = \left(Y_{kt} - \overline{Y}_k\right)^2$ represents the volatility of the log change in *Industry Sales*, for each 3-digit SIC sector *k*. $\overline{Y}_k$ is computed separately for the periods before and after bank entry deregulation.

[16] This measure is computed using the absolute value of the error term, after estimating equation (1) for the log changes in per capita income at the state level.

based on the use of public corporate debt. These variables have been widely used in the financing constraints literature (e.g., Almeida, Campello, and Weisbach (2004)). Empirical studies typically find that small firms and firms with limited access to corporate debt face significant information asymmetries when accessing credit markets,[17] and thus rely more on the use of internal funds or on credit from banks to finance their operations. Dependence on internal funds and bank credit makes these firms more likely to experience a change in financing conditions due to interstate bank entry deregulation. Formally, the following equation is estimated using volatility in production, sales, cash flow, and employment as dependent variables:

$$\begin{aligned} Vol\left(Y_{ijt}\right) = {} & \alpha_i + \mu_t + \delta Integration_{jt} + \lambda Bank\ Dependence_{ijt} \\ & + \phi Integration_{jt} \times Bank\ Dependence_{ijt} + \beta X_{ijt-1} + \gamma Vol(Z_{jt}) + \varphi Vol(W_{kt}) + \varepsilon_{ijt} \end{aligned} \tag{3}$$

where the right-hand side variables are the same as in equation (2), with the exception of *Bank Dependence* and its interaction with the dummy for bank integration. It is worth stressing that our two measures of *Bank Dependence*—the one based on firm size and the one based on access to public debt markets—are constructed so that larger values of the variable reflect a heavier dependence on banking finance. If bank-dependent firms benefit more from interstate deregulation, we would expect ϕ to have a negative sign. This implies that volatility decreases after deregulation for those firms that rely more heavily on banks for their external financing needs.

In the empirical estimations, the size-based measure of bank dependence is an indicator variable that equals one if the firm's assets are below median assets in a given year and state. Panel B in Table II shows that, according to this criterion, 42.1 percent of firm-year observations in our sample are classified as bank-dependent because of their small size. The second measure, based on the use of corporate debt, is defined through the firm's history of credit ratings and issues between 1970 and 1994. A firm is classified as being bank-dependent if it did not issue debt nor had any credit ratings in

[17] Gilchrist and Himmelberg (1995) use size to define financially constrained firms. These authors and Kashyap, Lamont, and Stein (1994) also use access to the bond market to define firms that face greater informational asymmetries in the financial markets. Calomiris, Himmelberg, and Wachtel (1995) describe the characteristics of commercial paper issuers.

this period. Bond and commercial paper credit rating information is retrieved from Compustat. Bond and commercial paper issues are obtained from the Mergent Fixed Income Security Database (FISD) and Moody's Default Risk Service (DRS) Database. Out of the 2,272 firms included in the sample, 1,516 are classified as bank-dependent because they did not have any issues or credit ratings for this period. As shown in Panel B of Table II, this group of firms accounts for 62.5 percent of firm-year observations.

D. Deregulation and the Cyclicality of Short-Term Borrowing

If bank lending availability is enhanced by deregulation, bank-dependent firms faced with idiosyncratic shocks would be able to borrow during periods of output contraction. Hence, bank credit becomes less pro-cyclical and bank entry deregulation would lead to smaller firm volatility. The next step in our analysis is to test whether bank credit became less pro-cyclical for bank-dependent firms after bank entry deregulation. For this purpose, we estimate the following equation:

$$\begin{aligned}\Delta \log\left(Notes\ Payable_{ijt}\right) = {} & \alpha_i + \mu_t + \delta Integration_{jt} + \theta Cyclicality_t + \\ & \lambda Bank\ Dependence_{ijt} + \phi Integration_{jt} \times Bank\ Dependence_{ijt} \\ & + \nu Integration_{jt} \times Bank\ Dependence_{ijt} \times Cyclicality_t \\ & + \beta X_{ijt-1} + \varepsilon_{ijt}\end{aligned} \quad (4)$$

where $\Delta \log\left(Notes\ Payable_{ijt}\right)$ is the log change in notes payable for firm *i* in state *j* and year *t*. Although loans from banks are not reported in Compustat, *Notes Payable* approximate access to credit from financial institutions.[18] We consider two alternative measures of the cycle: the first is firm sales scaled by lagged assets, $Sales_{ijt}/Assets_{ijt\text{-}1}$, and the second is the growth rate in real per capita GDP in the United States.

The coefficient on the cyclicality variables measures the co-movement of short-term borrowing with business conditions. In these estimations, we test if ν is negative and significant. When this coefficient is negative, cyclicality decreases after banking

[18] Notes payable is Compustat item #206.

deregulation for the set of firms that most likely use bank credit. This would be evidence that bank entry deregulation reduced the pro-cyclicality of lending to bank-dependent firms.

E. Equity Return Volatility

In addition to real variables, we study the effect of deregulation on equity return volatility. A change in firms' volatility, both of production and profitability, would likely have an effect on stock returns. As shown by Pastor and Veronesi (2003), idiosyncratic return volatility increases with the volatility of profitability. This is particularly relevant for financially constrained firms. A decrease in the volatility of financially constrained firms after bank entry deregulation should also reduce stock return volatility. To test this hypothesis, we measure stock return volatility in our baseline specifications as the standard deviation of the residuals from the following market model estimated with monthly observations:

$$\left(R_{ijt} - R_t^f\right) = \alpha_i + \beta_i \left(R_t^{vw} - R_t^f\right) + \varepsilon_{ijt} \tag{5}$$

where R_{ijt} is the return to shareholders of firm i in period t, which in our stock return regressions represents months instead of years. R_t^f is the risk-free return rate proxied by Ibbotson's one-month Treasury Bill rate; R_t^{VW} is the value-weighted return on all NYSE, NASDAQ, and AMEX stocks; and ε_{ijt} is the idiosyncratic component of excess returns. The market model in equation (5) is estimated for each firm before and after interstate entry deregulation, which allows for different coefficients in the two periods of interest. We also estimate (5) adding the Fama-French size and value factors, and the Jegadeesh-Titman momentum factor. Lastly, we analyze raw excess returns computed as the residuals in (5), without controlling for the excess returns on the market portfolio.

We define idiosyncratic stock return volatility as the standard deviation of residuals in (5) for each firm and year. Figures 1 and 2 show the evolution of these residual returns before and after deregulation. Following banking deregulation, the median idiosyncratic component of stock return volatility declines for three years. Campbell et al. (2001) and

Pastor and Veronesi (2003) document a long-term increase in firm volatility starting in the 1960s. This trend temporarily reverses in the years after deregulation in the 1980s, consistent with the drop in idiosyncratic stock return volatility in Figures 1 and 2.

[Figure 1]

[Figure 2]

Figure 3 shows the evolution of median stock return volatility before and after deregulation splitting the sample by bank dependence. Compared to the group of firms with access to public debt markets, excess return volatility is higher for firms that are bank-dependent and experiences a steeper decline after banking deregulation. This evidence suggests that the stabilizing effects from banking deregulation may have benefited more those firms that depend to a greater extent on banks for their financing needs. The next section presents the statistical analysis behind this graphical evidence.

[Figure 3]

III. Results

A. Firm Volatility

In this section we analyze the effect of interstate bank entry deregulation on firm volatility. The focus is on variables that account for firm output and performance. Table III presents the results of estimating equation (2) with the volatility of production, sales, cash flow, and employment as dependent variables. The coefficient on *Integration* is negative in all columns and statistically significant in the estimations for the volatility of cash flow and employment. To illustrate the economic magnitudes involved: after deregulation, the volatility of cash flow decreases by roughly 14 percent of the median volatility of cash flow in the sample. These results signal a sizeable decrease in volatility after out-of-state banks were permitted to enter local markets.

[Table III]

The findings in Table III also suggest that smaller and less profitable firms tend to be more volatile. Firm-specific volatility is not significantly correlated with state-level fluctuations. By contrast, industry-wide fluctuations appear to be an important component of firm volatility, as noted by the positive and significant coefficient on *Vol*(*Industry Sales*).

The results in this section are consistent with those in Morgan, Rime, and Strahan (2004), who find a decrease in the volatility of the growth of state-level employment due to banking deregulation. This decrease in volatility is associated with bank geographical diversification, and, therefore, less vulnerability to state-specific shocks. In the following sections, we study the mechanism that triggers the reduction in volatility at the firm level. But first we establish if bank-dependent firms are the ones that benefited the most from banking deregulation.

B. Firm Volatility and Bank Dependence

Small firms and firms without access to public debt markets are more likely to use bank credit to finance their operations. In the next set of tests, we compare the volatility of bank-dependent firms with the volatility of firms with access to other sources of finance, before and after the state deregulation of bank entry. These estimations allow us to control for changes in firm volatility within a state that are unrelated to decreases in out-of-state bank entry restrictions. Formally, we test whether volatility for bank-dependent firms decreased by more after interstate banking deregulation.

Table IV reports the results of estimating equation (3). In the regressions in Panel A, bank dependence is measured as lack of access to public debt markets, while, in Panel B, bank dependence is proxied by small firm size.[19] The coefficient on the interaction between bank dependence and the deregulation dummy is negative and significant in all specifications. This finding reflects the importance of deregulation on the observed decrease in volatility for the sample of bank-dependent firms. Bank deregulation makes

[19] Since the external finance proxy for debt issues and ratings is time-invariant, it drops from estimating equation (3) using firm-fixed effects.

bank-dependent firms more stable, but does not significantly alter the volatility of those firms that have ample access to non-bank sources of funding. The results in Table IV suggest that the effect of bank deregulation on firm volatility operates through direct relationship between firms and banks.

[Table IV]

The differential effect of deregulation on bank-dependent firms is substantial in most specifications, but it is particularly sizeable for cash-flow volatility. After deregulation, the reduction in the volatility of cash flow for small firms or for firms without access to public debt markets is about 27 percent relative to the median volatility of firm cash flow in the sample. The magnitude of the decrease in volatility of bank-dependent firms is somewhat smaller but still substantial when considering production, sales, and employment. As discussed by Booth and Cleary (2008), firms with more volatile cash flows hold larger amounts of financial slack to finance their investment. Lower volatility of output and cash flow and better access to external finance should decrease the need for cash holdings and increase their profitability.

The results outlined in this section confirm the significant effect of bank entry deregulation on the volatility of bank-dependent firms. After checking the robustness of our main results to different specifications, sample restrictions, and controls in the next section, we will assess whether the reduction in firm volatility is explained by a decrease in the pro-cyclicality of bank credit.

C. Robustness Checks

This section tests the sensitivity of our basic results in equations (2) and (3) to different specifications. In particular, Table V considers alternative volatility and deregulation measures, different sample selection criteria, and additional control variables. For the sake of compactness, Table V only reports the coefficients on the interstate deregulation dummy and its interaction with the bank dependence proxy (as measured by the firm's access to commercial paper and bond markets).[20]

[20] Results are very similar when we measure bank-dependence using a proxy related to firm size.

[Table V]

Panels A and B of Table V consider volatility measures alternative to the one defined in equation (1). Panel A reports estimates of regressions (2) and (3) computing the volatility of a generic variable Y_{ijt} as squared (instead of absolute) deviations of the variable from its conditional mean. Using the notation of equation (1), we use $\hat{\upsilon}_{ijt}^2$ instead of $\left|\hat{\upsilon}_{ijt}\right|$. Panel A suggests that giving a heavier weight to larger deviations in computing volatility does not affect qualitatively our basic results.

Panel B compares, for each firm, the five years before deregulation to the five years after deregulation, collapsing all observations into two periods. The dependent variables are the standard deviations of production, sales, cash flow, and employment in each period. Similarly, control variables are included in the regression as period averages. The results in Panel B show that the standard deviation of production, sales, and employment is lower in the five-year period after deregulation with respect to the five-year period after deregulation. The measure of volatility in Panel B is less noisy than the one we use in our baseline regressions, but it also makes hard to distinguish a causal effect of banking deregulation on firm volatility from an aggregate trend towards stability over the sample period.

Panel C replaces the banking deregulation dummy with a continuous measure of interstate banking integration: the share of deposits held by multi-state banks in each state and year. We instrument this continuous variable using the banking deregulation dummy. The results in Panel C indicate that firms become more stable when the fraction of deposits held by multi-state banks increases for the state where firms are located. In other words, *actual* integration of banks across states—the goal of deregulation—reduces volatility in non-financial firms, particularly for bank-dependent firms. The results in Panel C suggest that as the lending capacity of a state's "financial system" becomes more diversified through the participation of out-of-state banks, firms located in the state become more stable.

Panels D and E explore alternative sample selection criteria for regressions (2) and

(3). Some states deregulated intrastate banking around the same time they deregulated interstate banking. In order to isolate the effect interstate deregulation, Panel D excludes observations in states where intrastate banking was deregulated within a year of interstate bank-entry deregulation. Excluding these observations reduces the sample by about a sixth. Panel E exclude observations from the two states with the highest number of firms, California and New York, which account for about a quarter of the sample. The baseline results are insensitive to applying the more restrictive selection criteria of Panels D and E.

Panel F includes all available firm-year observations, unlike our baseline regressions, which include only firms with non-missing observations two years before and two years after deregulation. The results in Panel F are very similar to our baseline results. Hence, our finding that bank-dependent firms become more stable after deregulation is dominated by the decrease in volatility for surviving firms, leaving only a relatively small role for differences in across firms that entered after deregulation and those that exited before deregulation.

Panels G and H expand the set of control variables of the baseline regressions. In Panel G, we add a proxy for banking concentration to the right-hand side of equations (2) and (3). We measure banking concentration using the Herfindahl-Hirshman index of demand deposits by state. Panel H considers a larger set of firm-control variables. First, it includes leverage as measured by total debt scaled by assets. Second, it substitutes firm age, measured as time from IPO, for log sales. (Age and log-sales are not included simultaneously, as they are highly correlated.) The results from panel H suggest that the reduction in volatility after deregulation in previous sections is not explained by firms becoming more stable as they mature. Furthermore, in results not reported here, we estimate equations (2) and (3) replacing the actual banking deregulation dummy with a "placebo" deregulation dummy, constructed as an indicator that equals one starting two and four years after the actual deregulation took place in each state. We find that the effect of this "placebo" banking deregulation on the volatility of bank-dependent firms is not statistically significant.

The results summarized in Table V suggest that our main findings are robust to different volatility and deregulation measures, to alternative sample selection criteria, and

to the inclusion of additional controls. In practically all regressions reported in Table V, the interaction between the interstate deregulation dummy and the bank dependence proxy is negative and significant. This evidence suggests that firms that are more likely to rely on bank credit tend to be more stable, on average, after interstate banking deregulation.[21]

D. Cyclicality of Short-Term Borrowing

In this section we explore the channel that leads to the decrease in real volatility for our sample of publicly-traded firms. As Larrain (2006) points out, lower volatility has to be accompanied by increased countercyclical borrowing. As banks become more geographically diversified through deregulation, the correlation between bank capital and economic growth at the state level decreases (Morgan, Rime, and Strahan (2004)). Therefore, firms exploit bank credit to buffer from negative shocks.

We test the link between deregulation and firm volatility by checking the cyclicality of short-term borrowing. Table VI summarizes the results of estimating equation (4) using notes payable as the measure for short-term borrowing. Column (1) shows that notes payable is pro-cyclical on average. The coefficient on the interaction between the cyclicality variables and the deregulation dummy is negative but statistically insignificant. As expected, more profitable firms and with larger shares of tangible assets have higher borrowing growth rates.

[Table VI]

Columns (2) and (3) add proxies for bank dependence: public debt access and size. In both columns, the coefficient on the triple interaction between cyclicality, bank-dependence, and deregulation is negative and significant. Hence, short-term borrowing for small firms becomes less pro-cyclical after deregulation. The same pattern is observed for firms that do not use bonds or commercial paper to finance their operation.

[21] In other results not reported here, we replace for each state the actual deregulation date with a randomly generated date. The bank integration proxy and its interaction with the bank dependence measure become not significant.

This result is consistent with the decrease in the volatility of firm output and short-term credit described above. Geographically diversified banks are likely able to buffer firms from state-specific shocks by smoothing the credit they provide over the cycle.

E. Stock Return Volatility

In previous sections, we determined that output and cash flow volatility decreased after banking deregulation, especially for bank-dependent firms. This effect was accompanied by a reduction in the pro-cyclicality of credit. Based on our previous findings, we expect that lower volatility in real and financial variables after deregulation translates into less volatile stock returns. To test this hypothesis, we estimate the following equation:

$$\begin{aligned} Vol\left(R_{ijt}\right) = {} & \alpha_i + \mu_t + \delta Integration_{jt} + \lambda Bank\ Dependence_{ijt} \\ & + \phi Integration_{jt} \times Bank\ Dependence_{ijt} + \beta X_{ijt-1} + \varepsilon_{ijt} \end{aligned} \tag{6}$$

where $Vol(R_{ijt})$ is the standard deviation of idiosyncratic returns as defined in Section II.B. Our focus is on ϕ, the coefficient on the interaction between access to external finance and banking deregulation.

Table VII presents the result of estimating equation (6). The dependent variable in columns (1) through (3) is a measure of idiosyncratic stock return volatility computed as the standard deviations of excess returns on the firm stock (over the market return). The dependent variable in columns (4) through (6) measures idiosyncratic stock return volatility as the standard deviations of the residuals after estimating the market model for each firm. Idiosyncratic volatility in columns (7) through (8) is derived by adding the size and value factors from Fama and French, and the momentum factor from Jegadeesh and Titman (1993).

[Table VII]

In columns (1), (4), and (7) the coefficient on the deregulation indicator is negative but not statistically significant. On average, the decrease in idiosyncratic return volatility

is minimal. On the contrary, if we focus on bank-dependent firms, we find a considerable decrease in idiosyncratic return volatility after deregulation. These findings confirm the results in Pastor and Veronesi (2003) in a dynamic setting. As uncertainty about bank-dependent firms' future profitability decreases due to better access to external financing, stock return volatility for these firms also declines. In sum, the financing environment has a significant effect on a firm's real and financial volatility.

IV. Conclusion

The empirical methodology of this paper exploits the staggered timing of state deregulation of interstate banking in the United States in the 1980s and early 1990s. In previous decades, bank acquisition activity across state lines was generally very restricted. Deregulation of interstate banking promoted integration of state banking markets and enhanced banking competition. Part of the initial benefits of deregulation accrued to firms that relied on banks for funding, by improving the terms in which they were able to access banking finance. For example, Jayaratne and Strahan (1998), Dick (2006), and Rice and Strahan (2009) find that loan prices and spreads decrease after banking deregulation.

In our first set of results, we find that firms located in states that experienced interstate banking deregulation become more stable after deregulation. In particular, the growth rates of sales, production, cash flow, and employment become more stable for firms located in states that opened to interstate banking. Since the effect is stronger for those firms that depend on banks for external finance (because they are either small or have no credit ratings or issues), the stabilization is likely to be explained by the changes in the banking system induced by deregulation. When we exclude the proxies for bank dependence, the control group is the set of firms located in states that are yet to pass deregulation laws. When we include the proxies for bank dependence, we refine our control group to those firms that depend less on banks, because they can access external finance through public markets.

Our second set of results suggests that the stabilizing effects of banking deregulation are connected to the ability of firms to exploit credit to smooth out idiosyncratic shocks. In particular, short-term credit becomes more countercyclical after deregulation. Firms

may benefit from stability by smoothing investment when external financing is costly (Froot, Scharfstein, and Stein (1993)), reducing expected costs of financial distress, and lowering expected tax liabilities (Smith and Stulz (1985)). The concern for stability may be more pronounced for publicly traded firms, as Rountree, Weston, and Allayanis (2008) find that cash-flow volatility is negatively valued by investors.

In our third set of results, using a multifactor model of stock returns, we find that the idiosyncratic component of stock return volatility falls after deregulation. This finding suggests that the increased stability in operations (employment and production) and financial statements (cash flows) after deregulation translates into greater stock market stability. A reduction in idiosyncratic volatility may have an important impact on returns, because higher idiosyncratic volatility is associated with lower average returns (Ang et al. (2006a) and (2006b)).

Interpreting bank integration as a proxy for wider access to bank finance, the findings of this paper bridge two sets of results arguing that banking finance may have a stabilizing effect on the corporate sector. On the one hand, Larrain (2006) and Raddatz (2006) argue that industries that depend on banks are more stable in countries with more developed banking systems. On the other hand, Morgan, Rime, and Strahan (2004) show that banking deregulation had a stabilizing effect on state-level business cycles in the United States. Similar to Morgan, Rime, and Strahan (2004), we exploit interstate banking deregulation in the United States during the 1980s and early 1990s. This quasi-experiment addresses some potential concerns of using cross-country data. Similar to Larrain (2006) and Raddatz (2006), we exploit cross-sectional differences in the intensity of banking dependence. Unlike Larrain (2006), Raddatz (2006), and Morgan, Rime, and Strahan (2004), we are able to study the effect of banking development on individual stock returns, because we focus on firm-level data.

Recent research suggests that the reduction in macroeconomic volatility (the "Great Moderation" documented, for example, by McConnell and Perez-Quiroz (2000) and Stock and Watson (2002)) occurred despite an increase in volatility for publicly-traded firms over the second half of the 20th century.[22] Our results do not run against the long-

[22] In contrast to the publicly-traded firms' evidence, Davies et al. (2006), find a 23 percent decrease in employment growth volatility for privately held firms between 1978 and 2001.

term increase in firm-level volatility documented by Campbell et al. (2001), Comin and Philippon (2005), and Comin and Mulani (2006). In fact, the upward trend in volatility of those papers is temporarily halted or even reversed during the deregulation years we study in this paper. Our results do suggest, however, that the increase in firm volatility may have been steeper without interstate banking deregulation. In other words, firm-level volatility has increased despite of—not because of—banking deregulation. Naturally, the link between firm-level volatility and aggregate volatility deserves further attention.[23]

In our paper, the stabilizing effects of interstate banking deregulation were obtained starting from a situation where banking markets were highly fragmented across states. For those initial conditions, interstate banking deregulation likely generated efficiency gains that improved access to banking finance for the corporate sector. A deregulatory experiment starting from different conditions may not produce the same effects we find in this paper. Similarly, other forms of deregulation may trigger a different response on banks and, thus, on nonfinancial firms.

While the development of some financial institutions, like banks, may promote stability in the corporate sector, the development of other institutions, like equity markets, may have different effects. For example, Thesmar and Thoenig (2009) use a theoretical model to study the effect of international capital integration on the volatility of publicly-traded and privately-held firms. Similarly, using a panel of countries, Bekaert, Harvey, and Lundblad (2006) show that macroeconomic volatility falls after equity market liberalizations. Understanding the role of different institutions and their interaction is an interesting area for further research.[24]

[23] This paper complements existing evidence on the link between financial markets and financial innovation, on the one hand, and volatility, on the other. In the case of the United States, for example, Dynan, Elmendorf, and Sichel (2006) argue that financial innovation may help explain the stabilization of economic activity in the mid 1980s. An interesting avenue for exploration consists of studying the interaction between banking deregulation and financial innovation (in the form of more efficient pricing of risk, for example).

[24] See, for example, Acemoglu et al. (2003).

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Appendix 1: Dates of Deregulation of Interstate Bank Entry by State

State	*Year*	*State*	*Year*
ALASKA	1982	MONTANA	1993
ALABAMA	1987	NORTH CAROLINA	1985
ARKANSAS	1989	NORTH DAKOTA	1991
ARIZONA	1986	NEBRASKA	1990
CALIFORNIA	1987	NEW HAMPSHIRE	1987
COLORADO	1988	NEW JERSEY	1986
CONNECTICUT	1983	NEW MEXICO	1989
DISTRICT OF COLUMBIA	1985	NEVADA	1985
DELAWARE	1988	NEW YORK	1982
FLORIDA	1985	OHIO	1985
GEORGIA	1985	OKLAHOMA	1987
HAWAII	1995	OREGON	1986
IOWA	1991	PENNSYLVANIA	1986
IDAHO	1985	RHODE ISLAND	1984
ILLINOIS	1986	SOUTH CAROLINA	1986
INDIANA	1986	SOUTH DAKOTA	1988
KANSAS	1992	TENNESSEE	1985
KENTUCKY	1984	TEXAS	1987
LOUISIANA	1987	UTAH	1984
MASSACHUSETTS	1983	VIRGINIA	1985
MARYLAND	1985	VERMONT	1988
MAINE	1978	WASHINGTON	1987
MICHIGAN	1986	WISCONSIN	1987
MINNESOTA	1986	WEST VIRGINIA	1988
MISSOURI	1986	WYOMING	1987
MISSISSIPPI	1988		

Table I

Firm Distribution across States

The sample includes annual observations between 1976 and 1998 of Compustat firms classified in the manufacturing (SIC codes 2000 to 3999), wholesale (SIC codes 5000 to 5199), and retail trade (SIC codes 5200 to 5999) sectors. We consider only firms with non-missing information 2 years prior and 2 years after deregulation of interstate bank entry in the state where firms are headquartered.

State	*Number of firms*	*Pct. of total firms*	*Number of observations*	*Pct. of total observations*	*Average years per firm*
ALABAMA	13	0.6%	239	0.6%	18.4
ARKANSAS	8	0.4%	155	0.4%	19.4
ARIZONA	18	0.8%	284	0.7%	15.8
CALIFORNIA	312	13.7%	5,011	12.6%	16.1
COLORADO	41	1.8%	664	1.7%	16.2
CONNECTICUT	96	4.2%	1,585	4.0%	16.5
DISTRICT OF COLUMBIA	2	0.1%	46	0.1%	23.0
FLORIDA	84	3.7%	1,467	3.7%	17.5
GEORGIA	35	1.5%	661	1.7%	18.9
IOWA	14	0.6%	288	0.7%	20.6
IDAHO	3	0.1%	69	0.2%	23.0
ILLINOIS	137	6.0%	2,575	6.5%	18.8
INDIANA	42	1.8%	785	2.0%	18.7
KANSAS	13	0.6%	210	0.5%	16.2
KENTUCKY	12	0.5%	180	0.5%	15.0
LOUISIANA	8	0.4%	144	0.4%	18.0
MASSACHUSETTS	106	4.7%	1,809	4.6%	17.1
MARYLAND	34	1.5%	547	1.4%	16.1
MAINE	4	0.2%	59	0.1%	14.8
MICHIGAN	76	3.3%	1,449	3.7%	19.1
MINNESOTA	87	3.8%	1,574	4.0%	18.1
MISSOURI	46	2.0%	874	2.2%	19.0
MISSISSIPPI	2	0.1%	46	0.1%	23.0
MONTANA	1	0.0%	17	0.0%	17.0
NORTH CAROLINA	46	2.0%	845	2.1%	18.4
NEBRASKA	5	0.2%	98	0.2%	19.6
NEW HAMPSHIRE	9	0.4%	144	0.4%	16.0
NEW JERSEY	139	6.1%	2,379	6.0%	17.1
NEW MEXICO	2	0.1%	28	0.1%	14.0
NEVADA	5	0.2%	35	0.1%	7.0
NEW YORK	315	13.9%	5,104	12.9%	16.2
OHIO	122	5.4%	2,294	5.8%	18.8
OKLAHOMA	13	0.6%	244	0.6%	18.8
OREGON	15	0.7%	279	0.7%	18.6
PENNSYLVANIA	108	4.8%	2,065	5.2%	19.1
RHODE ISLAND	11	0.5%	183	0.5%	16.6
SOUTH CAROLINA	10	0.4%	179	0.5%	17.9
TENNESSEE	20	0.9%	380	1.0%	19.0
TEXAS	125	5.5%	2,179	5.5%	17.4
UTAH	15	0.7%	232	0.6%	15.5
VIRGINIA	43	1.9%	819	2.1%	19.0
VERMONT	3	0.1%	47	0.1%	15.7
WASHINGTON	31	1.4%	538	1.4%	17.4
WISCONSIN	39	1.7%	785	2.0%	20.1
WEST VIRGINIA	1	0.0%	15	0.0%	15.0
WYOMING	1	0.0%	14	0.0%	14.0
Total	2,272		39,624		17.4

Table II

Summary Statistics: Volatility and Access to External Finance

Panel A summarizes the properties of the firm-level volatility measures over the sample period, 1976-1998. *Vol*(*Y*) represents the volatility of *Y* as defined in Section II.B. *Production* is measured as *Sales* plus the change in inventories, and *Employment* is measured as total number of employees. These variables are taken from Compustat, and nominal variables are deflated using the GDP deflator. R_0 are firm-level excess returns over the Treasury Bill. R_1 is the residual of running a market-model regression of firm-level monthly excess returns on the excess return on the market portfolio and an intercept. R_2 is the residual obtained by expanding the market-model regression to include the size and value factors of Fama and French (1993) and the momentum factor of Jegadeesh and Titman (1993). Data on returns and factors are taken from CRSP and Kenneth French's Data Library at Dartmouth. Panel B summarizes firm-level proxies for *Bank Dependence*. The size-based measure of bank dependence equals 1 for firm-year observations when the firm assets are below the state-year median and 0 otherwise. The rating-issues-based measure equals 1 for firms that did not issue bonds or commercial paper and did not have credit ratings from 1970 to 1994. Data on bond and commercial paper issues are from Mergent Fixed Income Security Database and Moody's DRS. Ratings are taken from S&P as reported in Compustat.

Panel A

	Observations	*Mean*	*Median*	*Standard Deviation*
Vol (Production)	36,883	0.152	0.092	0.198
Vol (Sales)	36,905	0.132	0.081	0.159
Vol (Cash Flow)	29,679	0.353	0.196	0.489
Vol (Employment)	35,623	0.134	0.077	0.196
$Vol(R_0)$	28,270	0.119	0.104	0.070
$Vol(R_1)$	28,270	0.107	0.091	0.069
$Vol(R_2)$	28,270	0.103	0.087	0.066

Panel B

	Bank Dependence Proxied by Small Firm Size		*Bank Dependence Proxied by No Debt Issues and Ratings*	
	Observations	*Percentage*	*Observations*	*Percentage*
Not Dependent on Bank Finance	22,960	57.9%	14,849	37.5%
Dependent on Bank Finance	16,664	42.1%	24,775	62.5%

Table III

The Effect of Banking Deregulation on Firm-Level Volatility

This table reports the results from the regression:

$$Vol\left(Y_{ijt}\right) = \alpha_i + \mu_t + \delta Integration_{jt} + \beta X_{ijt-1} + \gamma Vol(Z_{jt}) + \varphi Vol(W_{kt}) + \varepsilon_{ijt}$$

Where *i* indexes firms, *j* denotes states, *k* indexes industries, and *t* represents years, and the sample period is 1976-1998. *Vol*(*Y*) is the volatility, as defined in Section II.B, of the growth rate of variable *Y*. We report results for 4 dependent variables: (1) *Vol*(*Production*), where *Production* is measured as *Sales* plus the change in inventories; (2) *Vol*(*Sales*); (3) *Vol*(*Cash Flow*); and (4) *Vol*(*Employment*), where *Employment* is measured as total number of employees. *Integration* is a dummy variable that equals 1 for state-year observations after laws that deregulated interstate banking were passed. *X* is a vector of 3 lagged firm controls: log(*Sales*); *EBITDA/Assets*, where EBITDA is earnings before income, tax, depreciation, and amortization; and *Tangible Assets/Assets*, where Tangible Assets are defined as in Almeida and Campello (2007). *Vol*(*Z*) is the volatility of state per capita income. *Vol*(*W*) is the volatility of industry sales, aggregated at the 3-digit SIC code level. Firm- and time-fixed effects are not reported. Firm-level variables are taken from Compustat, and nominal variables are deflated using the GDP deflator. State per capita income is from the Bureau of Economic Analysis. Robust standard errors clustered by state-year are reported in brackets below the coefficients. Significance: * significant at 10%; ** significant at 5%; *** significant at 1%.

Dependent variable:	***Vol(Production)***	***Vol(Sales)***	***Vol(Cash Flow)***	***Vol(Employment)***
	(1)	***(2)***	***(3)***	***(4)***
Integration	-0.004	-0.006	-0.027**	-0.007*
	[0.004]	[0.003]	[0.012]	[0.004]
log(*Sales* $_{t-1}$)	-0.021***	-0.021***	-0.004	-0.022***
	[0.003]	[0.003]	[0.007]	[0.003]
EBITDA $_{t-1}$ / *Assets* $_{t-1}$	-0.130***	-0.088***	-1.236***	-0.091***
	[0.025]	[0.014]	[0.064]	[0.018]
Tangible $_{t-1}$ / *Assets* $_{t-1}$	-0.011	-0.001	-0.194***	-0.006
	[0.021]	[0.016]	[0.048]	[0.027]
Vol(State p.c. Income)	-0.004	-0.030	-0.089	0.053
	[0.111]	[0.102]	[0.381]	[0.127]
Vol(Industry Sales)	0.109***	0.103***	0.087***	0.044***
	[0.012]	[0.009]	[0.029]	[0.010]
Observations	36,883	36,905	29,679	35,623
Firms	2,270	2,270	2,209	2,262
R-squared	0.02	0.03	0.04	0.01

Table IV

The Effect of Deregulation on Firm Volatility for Bank-Dependent Firms

This table reports the results from the regression:

$$Vol\left(Y_{ijt}\right) = \alpha_i + \mu_t + \delta Integration_{jt} + \lambda Bank\ Dependence_{ijt} + \phi Integration_{jt} \times Bank\ Dependence_{ijt} + \beta X_{ijt-1} + \gamma Vol(Z_{jt}) + \varphi Vol(W_{kt}) + \varepsilon_{ijt}$$

Where *i* indexes firms, *j* denotes states, *k* indexes industries, and *t* denotes years, and the sample period is 1976-1998. *Vol*(*Y*) is the volatility, as defined in Section II.B, of the growth rate of a generic variable *Y*. We report results for 4 dependent variables: (1) *Vol*(*Production*), where *Production* is measured as *Sales* plus the change in inventories; (2) *Vol*(*Sales*); (3) *Vol*(*Cash Flow*); and (4) *Vol*(*Employment*), where *Employment* is measured as total number of employees. *Integration* is a dummy variable that equals 1 for state-year observations after laws that deregulated interstate banking were passed. In Panel A, *Bank Dependence* equals 1 for firm-year observations when the firm assets are below the state-year median and 0 otherwise. In Panel B, *Bank Dependence* equals 1 for firms that did not issue bonds or commercial paper and did not have credit ratings from 1970 to 1994. *X* represents lagged firm controls: log(*Sales*); *EBITDA/Assets* and *Tangible Assets/Assets*. *Vol*(*Z*) is the volatility of state per capita income. *Vol*(*W*) is the volatility of industry sales, aggregated at the 3-digit SIC code level. Firm- and time-fixed effects are not reported. Firm-level variables are taken from Compustat, except for bond and commercial paper issues, which are from FISD and Moody's DRS. State per capita income is from the Bureau of Economic Analysis. Nominal variables are deflated using the GDP deflator. Robust standard errors clustered by state-year are reported in brackets below the coefficients. Significance: * significant at 10%; ** significant at 5%; *** significant at 1%.

PANEL A

	Bank Dependence Proxied by No Debt Issues and Ratings			
Dependent variable:	***Vol(Production)***	***Vol(Sales)***	***Vol(Cash Flow)***	***Vol(Employment)***
	(1)	***(2)***	***(3)***	***(4)***
Integration	0.004	0.003	0.016	0.002
	[0.005]	[0.004]	[0.013]	[0.005]
Bank Dependence	-	-	-	-
Integration ×	-0.013***	-0.013***	-0.069***	-0.013***
Bank Dependence	[0.004]	[0.003]	[0.011]	[0.004]
log*(Sales $_{t-1}$)*	-0.021***	-0.021***	-0.006	-0.023***
	[0.003]	[0.003]	[0.007]	[0.003]
EBITDA $_{t-1}$ / Assets $_{t-1}$	-0.130***	-0.088***	-1.245***	-0.091***
	[0.025]	[0.014]	[0.064]	[0.018]
Tangible $_{t-1}$ / Assets $_{t-1}$	-0.007	0.004	-0.157***	-0.001
	[0.021]	[0.016]	[0.050]	[0.027]
Vol(State p.c. Income)	0.000	-0.026	-0.080	0.057
	[0.112]	[0.102]	[0.377]	[0.128]
Vol(Industry Sales)	0.109***	0.103***	0.088***	0.044***
	[0.012]	[0.009]	[0.029]	[0.010]
Observations	36,883	36,905	29,679	35,623
Firms	2,270	2,270	2,209	2,262
R-squared	0.02	0.03	0.04	0.01

Table IV

The Effect of Deregulation on Firm Volatility for Bank-Dependent Firms (contd.)

PANEL B

	Bank Dependence Proxied by Small Firm Size			
Dependent variable:	***Vol(Production)***	***Vol(Sales)***	***Vol(Cash Flow)***	***Vol(Employment)***
	(1)	***(2)***	***(3)***	***(4)***
Integration	0.003	0.002	-0.003	-0.002
	[0.004]	[0.003]	[0.013]	[0.004]
Bank Dependence	-0.004	-0.001	0.043***	0.004
	[0.005]	[0.004]	[0.016]	[0.005]
Integration ×	-0.016***	-0.017***	-0.065***	-0.011**
Bank Dependence	[0.004]	[0.003]	[0.013]	[0.004]
log*(Sales $_{t-1}$)*	-0.023***	-0.023***	-0.003	-0.023***
	[0.004]	[0.003]	[0.008]	[0.003]
EBITDA $_{t-1}$ / Assets $_{t-1}$	-0.129***	-0.087***	-1.249***	-0.091***
	[0.025]	[0.014]	[0.064]	[0.018]
Tangible $_{t-1}$ / Assets $_{t-1}$	-0.006	0.005	-0.163***	-0.002
	[0.021]	[0.016]	[0.049]	[0.027]
Vol(State p.c. Income)	-0.006	-0.032	-0.099	0.052
	[0.111]	[0.101]	[0.379]	[0.127]
Vol(Industry Sales)	0.109***	0.103***	0.086***	0.044***
	[0.012]	[0.009]	[0.029]	[0.010]
Observations	36,883	36,905	29,679	35,623
Firms	2,270	2,270	2,209	2,262
R-squared	0.03	0.03	0.04	0.01

Table V

Robustness Checks

This table reports the results from the following two regressions:

$$Vol\left(Y_{ijt}\right) = \alpha_i + \mu_t + \delta Integration_{jt} + \beta X_{ijt-1} + \gamma Vol(Z_{jt}) + \varphi Vol(W_{kt}) + \varepsilon_{ijt}$$

$$Vol\left(Y_{ijt}\right) = \alpha_i + \mu_t + \delta Integration_{jt} + \lambda Bank\ Dependence_{ijt} + \phi Integration_{jt} \times Bank\ Dependence_{ijt} + \beta X_{ijt-1} + \gamma Vol(Z_{jt}) + \varphi Vol(W_{kt}) + \varepsilon_{ijt}$$

Where *i* indexes firms, *j* denotes states, *k* indexes industries, and *t* denotes years, and the sample period is 1976-1998. *Vol*(*Y*) is the volatility, as defined in Section II.B, of the growth rate of a generic variable *Y*. We report results for 4 dependent variables: (1) *Vol*(*Production*), where *Production* is measured as *Sales* plus the change in inventories; (2) *Vol*(*Sales*); (3) *Vol*(*Cash Flow*); and (4) *Vol*(*Employment*), where *Employment* is measured as total number of employees. α_i denotes firm-fixed effects and μ_t denotes year-fixed effects. All regressions include year-fixed effects, except for Panel B. *Integration* is a dummy variable that equals 1 for state-year observations after laws that deregulated interstate banking were passed. *Bank Dependence* equals 1 for firms that did not issue bonds or commercial paper and did not have credit ratings from 1970 to 1994. Data on debt issues are from FISD and Moody's DRS. *X* represents lagged firm controls: log(*Sales*); *EBITDA/Assets* and *Tangible Assets/Assets*. *Vol*(*Z*) is the volatility of state per capita income. *Vol*(*W*) is the volatility of industry sales, aggregated at the 3-digit SIC code level. Firm-level variables are taken from Compustat, except for bond and commercial paper issues, which are from FISD and Moody's DRS. State per capita income is from the Bureau of Economic Analysis. Nominal variables are deflated using the GDP deflator. Robust standard errors clustered by state-year are reported in brackets below the coefficients. Significance: * significant at 10%; ** significant at 5%; *** significant at 1%.

Panel A - Dependent variable is volatility, measured by the squared value of deviations from the firm-level conditional mean

	Bank Dependence Proxied by No Debt Issues and Ratings							
Dependent var.:	***Vol(Production)***		***Vol(Sales)***		***Vol(Cash Flow)***		***Vol(Employment)***	
	(1)	***(2)***	***(3)***	***(4)***	***(5)***	***(6)***	***(7)***	***(8)***
Integration	0.003	0.010	-0.005	0.003	-0.062*	0.002	-0.006	0.005
	[0.008]	[0.008]	[0.003]	[0.003]	[0.033]	[0.037]	[0.007]	[0.010]
Integration ×		-0.011*		-0.011***		-0.104***		-0.016*
Bank Dependence		[0.006]		[0.003]		[0.033]		[0.010]
Observations	36,883	36,883	36,905	36,905	29,679	29,679	35,623	35,623
Firms	2,270	2,270	2,270	2,270	2,209	2,209	2,262	2,262
R-squared	0.01	0.01	0.02	0.02	0.02	0.02	0.00	0.00

Panel B - Dependent variable is volatility, measured as the standard deviation five years before and five years after banking deregulation

	Bank Dependence Proxied by No Debt Issues and Ratings							
Dependent var.:	***Vol(Production)***		***Vol(Sales)***		***Vol(Cash Flow)***		***Vol(Employment)***	
	(1)	***(2)***	***(3)***	***(4)***	***(5)***	***(6)***	***(7)***	***(8)***
Integration	-0.007	0.005	-0.005	0.010*	0.051***	0.048***	0.004	0.015
	[0.007]	[0.007]	[0.006]	[0.005]	[0.013]	[0.015]	[0.007]	[0.009]
Integration ×		-0.017***		-0.024***		0.005		-0.017**
Bank Dependence		[0.006]		[0.005]		[0.023]		[0.008]
Observations	2,929	2,929	2,929	2,929	2,649	2,649	2,868	2,868
Firms	1,469	1,469	1,469	1,469	1,420	1,420	1,454	1,454
R-squared	0.03	0.03	0.04	0.05	0.08	0.08	0.04	0.04

Table V

Robustness Checks (contd.)

Panel C - Bank integration is measured as the share of deposits held by multi-state banks in each state and year (This measure is instrumented by the deregulation dummy)

	Bank Dependence Proxied by No Debt Issues and Ratings							
Dependent var.:	***Vol(Production)***		***Vol(Sales)***		***Vol(Cash Flow)***		***Vol(Employment)***	
	(1)	***(2)***	***(3)***	***(4)***	***(5)***	***(6)***	***(7)***	***(8)***
Integration	-0.040	-0.018	-0.051	-0.030	-0.232*	-0.134	-0.063	-0.041
	[0.044]	[0.043]	[0.032]	[0.031]	[0.120]	[0.118]	[0.040]	[0.039]
Integration ×		-0.030***		-0.030***		-0.158***		-0.031***
Bank Dependence		[0.010]		[0.008]		[0.031]		[0.010]
Observations	36,881	36,881	36,903	36,903	29,634	29,634	35,614	35,614
Firms	2,268	2,268	2,268	2,268	2,164	2,164	2,253	2,253
R-squared	0.02	0.02	0.02	0.02	0.03	0.04	0.01	0.01

Panel D - Sample excludes states with simultaneous intrastate and interstate banking deregulation

	Bank Dependence Proxied by No Debt Issues and Ratings							
Dependent var.:	***Vol(Production)***		***Vol(Sales)***		***Vol(Cash Flow)***		***Vol(Employment)***	
	(1)	***(2)***	***(3)***	***(4)***	***(5)***	***(6)***	***(7)***	***(8)***
Integration	-0.006	0.004	-0.006	0.004	-0.024**	0.014	-0.007	0.004
	[0.004]	[0.005]	[0.004]	[0.004]	[0.012]	[0.014]	[0.004]	[0.006]
Integration ×		-0.014***		-0.014***		-0.062***		-0.017***
Bank Dependence		[0.004]		[0.004]		[0.012]		[0.004]
Observations	30,652	30,652	30,672	30,672	24,605	24,605	29,636	29,636
Firms	1,896	1,896	1,896	1,896	1,843	1,843	1,890	1,890
R-squared	0.03	0.03	0.03	0.03	0.04	0.04	0.01	0.02

Panel E - Sample excludes firms in California and New York

	Bank Dependence Proxied by No Debt Issues and Ratings							
Dependent var.:	***Vol(Production)***		***Vol(Sales)***		***Vol(Cash Flow)***		***Vol(Employment)***	
	(1)	***(2)***	***(3)***	***(4)***	***(5)***	***(6)***	***(7)***	***(8)***
Integration	0.002	0.012*	-0.001	0.009*	-0.020	0.036**	-0.008	0.001
	[0.006]	[0.006]	[0.005]	[0.005]	[0.015]	[0.016]	[0.006]	[0.006]
Integration ×		-0.014***		-0.015***		-0.089***		-0.014***
Bank Dependence		[0.004]		[0.003]		[0.012]		[0.005]
Observations	27,540	27,540	27,558	27,558	22,690	22,690	26,565	26,565
Firms	1,645	1,645	1,645	1,645	1,608	1,608	1,639	1,639
R-squared	0.02	0.02	0.02	0.02	0.04	0.04	0.01	0.01

Panel F - Sample includes all firms with non-missing observations between 1976 and 1998

	Bank Dependence Proxied by No Debt Issues and Ratings							
Dependent var.:	***Vol(Production)***		***Vol(Sales)***		***Vol(Cash Flow)***		***Vol(Employment)***	
	(1)	***(2)***	***(3)***	***(4)***	***(5)***	***(6)***	***(7)***	***(8)***
Integration	-0.002	0.006	-0.002	0.006*	-0.030***	0.009	-0.005	0.004
	[0.004]	[0.004]	[0.004]	[0.004]	[0.011]	[0.013]	[0.004]	[0.005]
Integration ×		-0.013***		-0.012***		-0.062***		-0.014***
Bank Dependence		[0.004]		[0.003]		[0.010]		[0.004]
Observations	60,939	60,939	60,987	60,987	45,723	45,723	58,442	58,442
Firms	7,288	7,288	7,290	7,290	6,046	6,046	7,104	7,104
R-squared	0.02	0.02	0.02	0.02	0.03	0.04	0.01	0.01

Table V

Robustness Checks (contd.)

Panel G - Controls include Banking Concentration as measured by Herfindahl-Hirshman index of deposits by state

	Bank Dependence Proxied by No Debt Issues and Ratings							
Dependent var.:	***Vol(Production)***		***Vol(Sales)***		***Vol(Cash Flow)***		***Vol(Employment)***	
	(1)	***(2)***	***(3)***	***(4)***	***(5)***	***(6)***	***(7)***	***(8)***
Integration	-0.004	0.004	-0.006*	0.003	-0.027**	0.016	-0.007*	0.002
	[0.004]	[0.005]	[0.003]	[0.004]	[0.012]	[0.013]	[0.004]	[0.005]
Integration ×		-0.013***		-0.013***		-0.069***		-0.013***
Bank Dependence		[0.004]		[0.003]		[0.011]		[0.004]
Observations	36,883	36,883	36,905	36,905	29,679	29,679	35,623	35,623
Firms	2,270	2,270	2,270	2,270	2,209	2,209	2,262	2,262
R-squared	0.02	0.02	0.03	0.03	0.04	0.04	0.01	0.01

Panel H - Controls include age (as measured by years from IPO) instead of log(Sales), and firm leverage, measured as total debt divided by total assets

	Bank Dependence Proxied by No Debt Issues and Ratings							
Dependent var.:	***Vol(Production)***		***Vol(Sales)***		***Vol(Cash Flow)***		***Vol(Employment)***	
	(1)	***(2)***	***(3)***	***(4)***	***(5)***	***(6)***	***(7)***	***(8)***
Integration	-0.008*	-0.001	-0.008**	-0.003	-0.037***	-0.009	-0.006	-0.002
	[0.005]	[0.005]	[0.003]	[0.004]	[0.013]	[0.015]	[0.004]	[0.005]
Integration ×		-0.011**		-0.009***		-0.049***		-0.006
Bank Dependence		[0.004]		[0.003]		[0.012]		[0.004]
Observations	30,592	30,592	30,604	30,604	25,221	25,221	30,025	30,025
Firms	2,015	2,015	2,015	2,015	1,939	1,939	2,007	2,007
R-squared	0.02	0.02	0.02	0.02	0.04	0.04	0.01	0.01

Table VI

Banking Deregulation and Countercyclical Borrowing

This table reports the results from the regression:

$$\Delta \log\left(Notes\ Payable_{ijt}\right) = \alpha_i + \mu_t + \delta Integration_{jt} + \theta Cyclicality_t + \lambda Bank\ Dependence_{ijt} + \phi Integration_{jt} \times Bank\ Dependence_{ijt} + \nu Integration_{jt} \times Bank\ Dependence_{ijt} \times Cyclicality_t + \beta X_{ijt-1} + \varepsilon_{ijt}$$

Where *i* indexes firms, *j* denotes states, *k* indexes industries, and *t* denotes years, and the sample period is 1976-1998. The dependent variable is the log change in *Notes Payable*. Estimations in Columns (4) to (6) do not include year-fixed effects. *Integration* is a dummy variable that equals 1 for state-year observations after interstate banking deregulation laws were passed. The table reports results with and without two proxies for bank dependence. The size-based measure of *Bank Dependence* equals 1 for firm-year observations when the firm assets are below the state-year median and 0 otherwise. The rating-issues-based measure of *Bank Dependence* equals 1 for firms that did not issue bonds or commercial paper and did not have credit ratings from 1970 to 1994. *X* represents lagged firm controls: log(*Sales*); *EBITDA/Assets* and *Tangible Assets/Assets*. Firm-level variables are taken from Compustat, except for bond and commercial paper issues, which are from FISD and Moody's DRS. Nominal variables are deflated using the GDP deflator. Robust standard errors clustered by state-year are reported in brackets below the coefficients. Significance: * significant at 10%; ** significant at 5%; *** significant at 1%.

Dependent variable:	***Δlog (Notes Payable)***					
Cyclicality measured as:	***Sales$_t$ / Assets$_{t-1}$***			**Δlog (*Real GDP*)**		
Bank dependence measured as:	***No Bank Depend. Proxy***	***No Debt Issues and Ratings***	***Small Firm Size***	***No Bank Depend. Proxy***	***No Debt Issues and Ratings***	***Small Firm Size***
	(1)	***(2)***	***(3)***	***(4)***	***(5)***	***(6)***
Integration	-0.007	-0.027	-0.032	0.046	0.046	0.046
	[0.061]	[0.059]	[0.058]	[0.036]	[0.036]	[0.036]
Cyclicality	0.127***	0.139***	0.148***	3.837***	3.825***	3.819***
	[0.029]	[0.027]	[0.026]	[0.533]	[0.532]	[0.531]
Bank Dependence			0.114**			0.092*
			[0.055]			[0.053]
Integration × *Cyclicality*	-0.019	0.056*	0.047	-1.577	-0.275	-0.257
	[0.028]	[0.030]	[0.030]	[0.966]	[1.127]	[1.051]
Integration × *Bank Dependence* × *Cyclicality*		-0.100***	-0.125***		-2.395**	-3.941***
		[0.024]	[0.029]		[0.954]	[0.975]
log(*Sales $_{t-1}$*)	-0.044*	-0.053**	-0.047*	-0.059**	-0.064***	-0.059**
	[0.024]	[0.024]	[0.025]	[0.024]	[0.024]	[0.025]
EBITDA $_{t-1}$ / Assets $_{t-1}$	0.678***	0.672***	0.670***	0.768***	0.770***	0.770***
	[0.099]	[0.098]	[0.098]	[0.104]	[0.104]	[0.104]
Tangible $_{t-1}$ / Assets $_{t-1}$	0.507***	0.542***	0.532***	0.624***	0.666***	0.678***
	[0.194]	[0.194]	[0.191]	[0.179]	[0.181]	[0.180]
Observations	16,953	16,953	16,953	16,953	16,953	16,953
Firm	1,875	1,875	1,875	1,875	1,875	1,875
R-squared	0.02	0.02	0.02	0.01	0.01	0.01

Table VII

Banking Deregulation and Idiosyncratic Stock Return Volatility

This table presents the results from the following regression:

$$Vol\left(R_{it}\right) = \alpha_i + \mu_t + \delta Integration_{jt} + \lambda Bank\ Dependence_{ijt} + \phi Integration_{jt} \times Bank\ Dependence_{ijt} + \beta X_{ijt-1} + \varepsilon_{ijt}$$

Where i indexes firms, j denotes states, and t denotes year, and the sample period is 1976-1998. R_0 are firm-level raw excess returns over the Treasury Bill. R_1 is the residual of running a firm-level market-model regression of monthly excess returns on the excess return on the market portfolio and an intercept. R_2 is the residual obtained by expanding the market model regression to include the size and value factors of Fama and French (1993) and the momentum factor of Jegadeesh and Titman (1993). $Vol(R_{it})$ is the volatility of R_{it}, measured as its standard deviation in year t. *Integration* is a dummy variable that equals 1 for state-year observations after interstate banking deregulation laws were passed. The size-based measure of *Bank Dependence* equals 1 for firm-year observations when the firm assets are below the state-year median and 0 otherwise. The rating-issues-based measure of *Bank Dependence* equals 1 for firms that did not issue bonds and commercial paper and did not have credit ratings from 1970 to 1994. X represents lagged firm controls: log(*Sales*); *EBITDA/Assets* and *Tangible Assets/Assets*. Firm-level variables are taken from Compustat, except for bond and commercial paper issues, which are from FISD and Moody's DRS. Nominal variables are deflated using the GDP deflator. Robust standard errors clustered by state-year are reported in brackets below the coefficients. Significance: * significant at 10%; ** significant at 5%; *** significant at 1%.

Dependent variable:	***Vol (R_0)***			***Vol (R_1)***			***Vol (R_2)***		
Bank dependence measured as:	***No Bank Depend. Proxy***	***No Debt Issues and Ratings***	***Small Firm Size***	***No Bank Depend. Proxy***	***No Debt Issues and Ratings***	***Small Firm Size***	***No Bank Depend. Proxy***	***No Debt Issues and Ratings***	***Small Firm Size***
	(1)	***(2)***	***(3)***	***(4)***	***(5)***	***(6)***	***(7)***	***(8)***	***(9)***
Integration	-0.001	0.005***	0.002	-0.001	0.003*	0.001	-0.001	0.002	0.001
	[0.001]	[0.002]	[0.001]	[0.001]	[0.002]	[0.001]	[0.001]	[0.002]	[0.001]
Bank Dependence			0.011***			0.011***			0.010***
			[0.002]			[0.002]			[0.002]
Integration × *Bank Dependence*		-0.008***	-0.006***		-0.006***	-0.006***		-0.005***	-0.005**
		[0.002]	[0.002]		[0.001]	[0.002]		[0.001]	[0.002]
log(*Sales* $_{t-1}$)	-0.013***	-0.013***	-0.011***	-0.013***	-0.013***	-0.012***	-0.013***	-0.013***	-0.011***
	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]
EBITDA $_{t-1}$ / *Assets* $_{t-1}$	-0.071***	-0.071***	-0.073***	-0.081***	-0.081***	-0.083***	-0.076***	-0.076***	-0.078***
	[0.009]	[0.009]	[0.009]	[0.008]	[0.008]	[0.008]	[0.008]	[0.008]	[0.008]
Tangible $_{t-1}$ / *Assets* $_{t-1}$	-0.047***	-0.044***	-0.045***	-0.049***	-0.047***	-0.047***	-0.045***	-0.043***	-0.044***
	[0.007]	[0.007]	[0.007]	[0.007]	[0.007]	[0.007]	[0.007]	[0.007]	[0.007]
Observations	28,167	28,167	28,167	28,167	28,167	28,167	28,167	28,167	28,167
Firms	1,637	1,637	1,637	1,637	1,637	1,637	1,637	1,637	1,637
R-squared	0.09	0.10	0.10	0.07	0.07	0.07	0.06	0.06	0.06

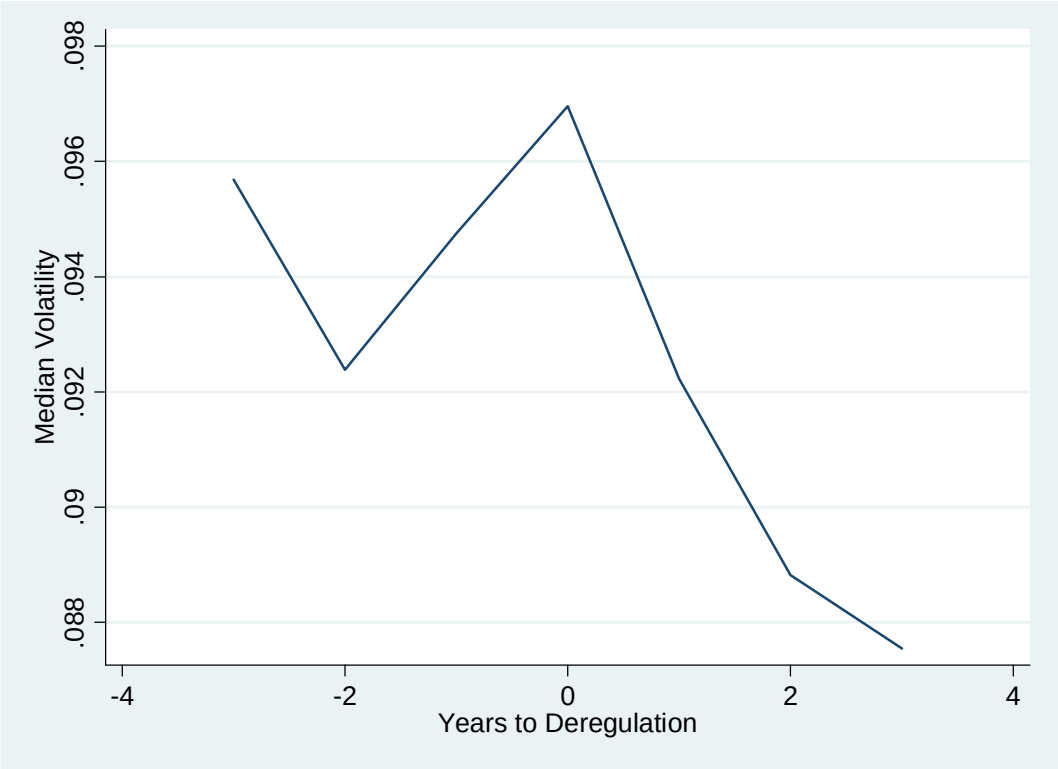

Figure 1. Volatility of idiosyncratic returns estimated from a market model. The figure shows the median idiosyncratic component of firm volatility for the three years before and after interstate bank entry deregulation. Volatility is measured as the standard deviation of residuals from a market model.

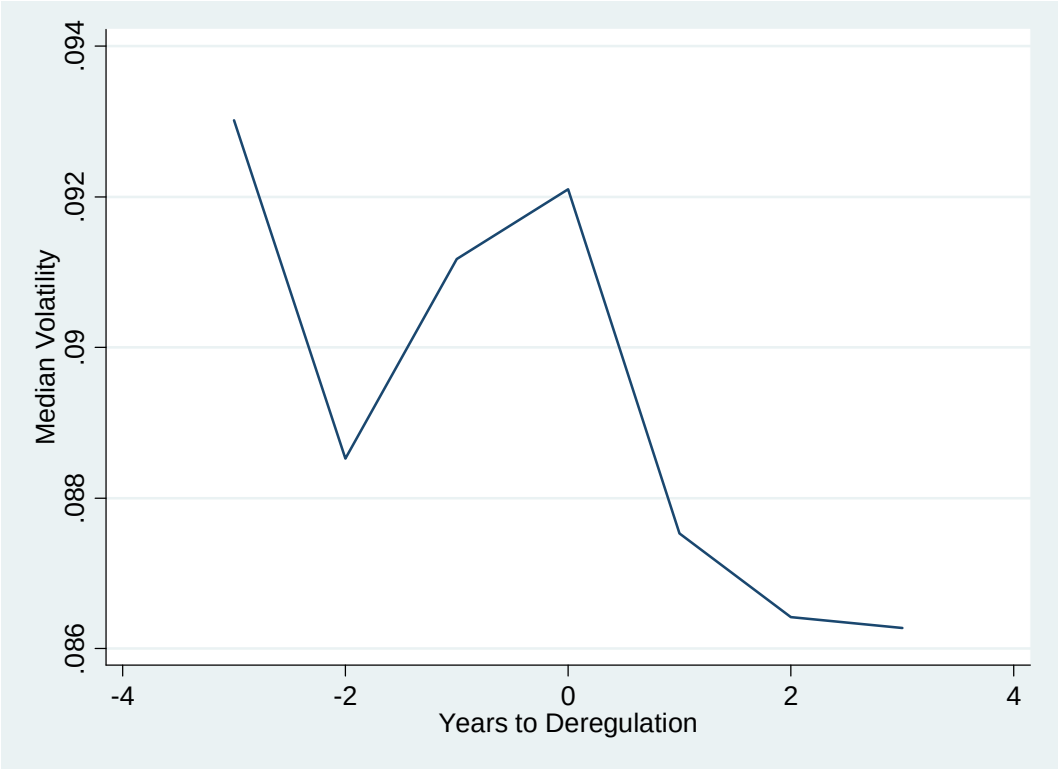

Figure 2. Volatility of idiosyncratic returns estimated from a four-factor model. The figure shows the median idiosyncratic component of firm volatility for the three years before and after interstate bank entry deregulation. Volatility is measured as the standard deviation of residuals from a four-factor model.

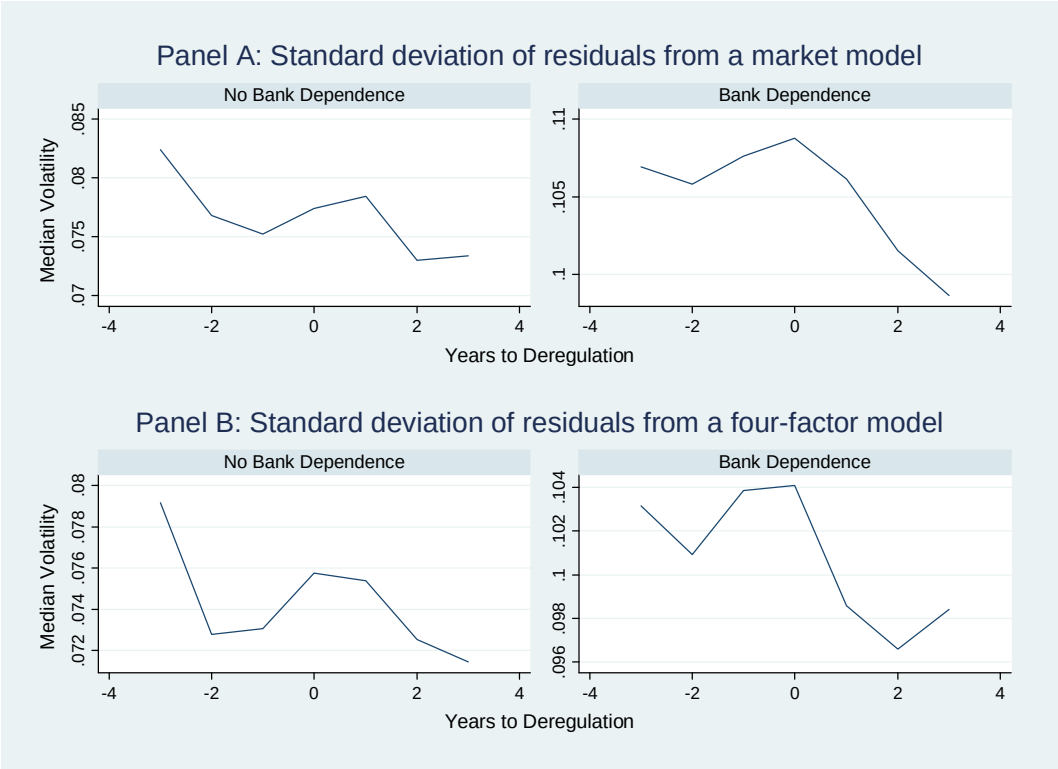

Figure 3. Volatility of idiosyncratic returns and bank dependence. The figure shows the median idiosyncratic component of firm volatility for the three years before and three years after interstate bank entry deregulation. Volatility is measured as the standard deviation of residuals from a market model (Panel A) and a four-factor model (Panel B). The sample of firms is divided according to bank dependence. A firm is classified as being bank dependent if it did not access public debt markets between 1970 and 1994.