The Impact of Banking Deregulation on Inbound Foreign Direct Investment: Transaction-level Evidence from the United States*

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Abstract:

We evaluate the effects of state-level banking deregulation that resulted in improved access to cheaper local finance on foreign firms investing in the U.S. We provide direct, micro-level evidence from U.S. inbound foreign direct investment transactions showing that interstate banking, but not intrastate branching, deregulation increased the number of transactions, reduced the average transaction value, and boosted overall investment by foreign multinationals. We also show that lower cost of local credit and greater local bank competition in each state, following the interstate banking deregulation, are potential mechanisms that stimulated FDI activity. Finally, we demonstrate that after the adoption of the interstate banking deregulation, both the number and the average value of transactions increased in industries that are more dependent on external finance relative to industries that are less dependent.

Keywords: Foreign Direct Investment; Banking Deregulation; External Finance Dependence
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1. Introduction

Until the early 1970s most U.S. states either prohibited or severely restricted both interstate banking and intrastate bank branching. In the late 1970s, many states began lifting restrictions on intrastate bank branching and interstate bank expansions. These two types of deregulation led to higher competition, greater efficiency, and reduction in monopoly power in the banking sector, thereby facilitating access to cheaper local credit (Jayaratne & Strahan 1996; Jayaratne & Strahan 1998; Cetorelli & Strahan 2006).[1] A number of studies have examined the subsequent effects on domestic U.S. firms in the financial and manufacturing sectors. However, no work has been done to date to evaluate the impact of the two banking deregulations and the accompanying reduction in the cost of credit on foreign firms entering the U.S. market. This study attempts to fill this gap by providing direct, micro-level evidence from U.S. inbound foreign direct investment (FDI) transactions.[2]

Our main hypothesis is that the banking deregulations had a positive impact on FDI activity. We know from the existing literature that multinational firms utilize significant amounts of host country debt financing in their affiliates' capital structure.[3] Such financing is used both for cross-border transactions as

[1] Strahan (2003) argues that banking deregulation has resulted in larger banks operating across broader geographic areas, but has not brought about higher concentrations at the local level. Banks also became more efficient: for instance, Jayaratne & Strahan (1998) find that in the long run, costs to borrowers decrease by 0.3 percent, loan losses decrease by half a percent, and operating costs decline by 8 percent.

[2] Throughout this paper, we will use the term FDI to refer to *inbound* FDI into the U.S. *Outbound* FDI, originating from the U.S. and flowing to other countries is outside the scope of our study.

[3] Host country borrowing by multinationals was prevalent throughout our sample period (1977-1994). Horst (1977) estimates that of the $21 billion of foreign investment made by U.S. multinationals in 1974, some $18.3 billion was financed through host country debt as well as retained earnings. Examining data from the end of our sample period, Feldstein (1995) reports that U.S. investment in non-bank controlled foreign corporations in 1989 totaled $1,237 billion, of which $567 billion was financed through non-U.S. debt. Using a comprehensive dataset of all foreign affiliates of U.S. multinationals, Desai *et al.* (2004) estimate that foreign affiliates had an external borrowing to assets ratio of over 44 percent. The same pattern holds true for the U.S. affiliates of foreign multinationals. Laster and McCauley (1994) document that between 1979 and 1992 the leverage ratio, excluding intercompany debt (i.e. excluding debt from parent firms) for foreign firms operating in the U.S. averages 44 percent, the majority of which is financed in the U.S. Once intercompany debt is included, the leverage ratio of foreign affiliates rises to 57 percent, suggesting that external host country borrowing is a more important source of debt financing than intrafirm borrowing. Similarly, Marin and Schnitzer (2011) provide evidence that Eastern European affiliates of German and Austrian firms source 30 to 40 percent of their external financing needs from local sources.

well as for the ongoing operations of foreign affiliates.[4] Therefore, variation in the cost of external local debt finance could play a significant role for the incidence and the intensity of cross-border transactions.

To assess the importance of cheaper local credit on inbound FDI, we estimate the impact of U.S. interstate banking and intrastate branching deregulations on the number and the size of inbound FDI transactions. We show that the interstate banking deregulation was associated with a higher entry rate of foreign multinationals, a larger number of FDI transactions, and a smaller average transaction value while the deregulation of intrastate branching did not have any significant effect on FDI inflows along the extensive or the intensive margin.[5] We also find that as the fraction of states that allow interstate banking grew, the overall volume of inbound FDI undertaken by foreign multinationals increased. In particular, our empirical evidence suggests that on average a state, which adopted the interstate banking deregulation experienced a 19 percent increase in the number of inbound FDI transactions, translating to 1.28 new transactions per year, and an increase in the entry rate of foreign multinationals of about 42 percent.

Investigating the impact of banking deregulation along the intensive margin, we find that the average value of foreign transactions decreased by approximately 27.4 percent following the adoption of the interstate banking deregulation. The result is robust to including a comprehensive list of state-level, time-varying controls and trends, as well as source country and mode of entry fixed effects. Our results indicate that with cheaper external finance, foreign firms were able to undertake projects of smaller value,

[4] Faccio and Masulis (2005) show that cross-border merger and acquisition deals are more likely to be financed with cash as opposed to stock, and cash transactions in turn are likely to involve external borrowing. Beyond the initial transaction, debt is also extensively used to finance the continued operations of foreign affiliates, which typically use a mix of internal and external host country debt financing. The use of host country financing as a means to manage tax liabilities has been discussed at length in the international tax context (see, for example, Gresik (2001) and Graham (2003)). Chowdhry and Nanda (1994) present a theoretical model in which parent firms finance their foreign affiliates with a combination of internal and external debt, taking advantage of the tax advantaged nature of debt. In their model, external local debt serves as a benchmark for setting the rate for internal borrowing. Host country financing is also an effective means of hedging against currency risk (Graham and Harvey (2001)). External local debt financing is more widely used in countries with lower political risk (Desai *et al.* (2008)). Desai *et al.* (2004) show that external local debt financing is particularly popular in countries with well-developed capital markets and strong creditor rights, such as the U.S., because the cost of borrowing is lower.

[5] We define the entry rate as the ratio of new FDI transactions to the total number of existing multinationals in a given state and year, i.e. as the share of new transactions. The extensive margin refers to the incidence of FDI or the number of transactions and the entry rate. It captures the *gross* entry rate, as the ITA data provides information only on the new FDI transactions undertaken by foreign multinationals, and it does not report the multinational firms that exit. The intensive margin, on the other hand, refers to the intensity of FDI activity or transaction values.

which became more profitable when borrowing costs declined. Further, we demonstrate that when the share of states, which allow interstate banking, rose, overall investment in the U.S. undertaken by foreign multinationals grew. Our estimates suggest that as the share increased by 10 percent (equivalent to 5 additional states adopting the interstate banking deregulation), foreign firms' overall investment in the U.S. rose by 14.4 percent, which corresponds to an increase in total FDI inflows into the U.S. manufacturing sector of 1.9 billion (1983 U.S. $), or 8.2 percent of the total FDI inflows.

To illuminate the mechanisms behind the effect of banking deregulation on the incidence and the intensity of FDI, we extend our work in two directions. First, we consider how FDI responded to changes in the cost of credit and bank industry structure resulting from the banking deregulations. We show that lower cost of credit and greater bank competition stimulated FDI activity.[6] Second, we provide direct evidence of the importance of the local finance channel for FDI by comparing the impact of banking deregulation on foreign transactions taking place in sectors that rely on external finance more heavily versus those in sectors that are less reliant on external finance (Rajan & Zingales 1998; Cetorelli & Strahan 2006). If access to cheaper, local finance were important for inbound FDI activity, one would anticipate the effects of banking deregulation to be more pronounced in industries that are more reliant on external finance.

Consistent with prior studies, we confirm that interstate banking deregulation significantly lowered the cost of credit as measured by the loan yield, greatly enhanced competition in the banking industry as measured by the Herfindahl-Hirschmann index (HHI) while at the same time increasing the share of assets held by large banks in deregulated states. We find that these structural changes in the bank industry have a sizeable effect on FDI activity. Using the interstate banking deregulation as an instrument

[6] While it is impossible to provide direct evidence, it may also be the case that large national banks have a comparative advantage in evaluating and financing FDI projects and hence interstate banking deregulation, which led to the advent of national banks, would encourage greater FDI activity through this channel. We find some suggestive evidence showing that the impact of interstate banking on multi-state foreign investors is stronger. This could be because economies of scale can emerge from the opportunity for a foreign investor to exploit a relationship with a single, large, national bank after interstate banking deregulation (as opposed to multiple, smaller, local financial institutions). Further, multi-state investors may be more likely to avail themselves of local bank finance since they have prior exposure to the U.S. market (and a higher likelihood of local collateral), which one-time, single-state investors lack.

for changes in the state banking environment, we find that lower loan yields are associated with greater foreign entry and lower transaction values. Similarly, greater bank competition, measured by a lower HHI, leads to greater foreign entry and lower transaction value. Finally, higher share of large bank assets is also associated with greater foreign entry and lower average transaction values. This evidence identifies the cost of credit and the banking industry structure as direct mechanisms behind the effect of banking deregulation on inbound FDI across U.S. states.

Turning to the effect of the interstate banking deregulation on FDI activity in sectors that are *more* dependent on external finance, we find that following the adoption of the deregulation, the increase in the entry rate of foreign multinationals was far more pronounced in industries that are *more* dependent on external finance. Hence, by facilitating access to credit, interstate banking deregulation allowed a larger number of foreign firms that rely on external finance more heavily to invest in the U.S. Along the intensive margin we find that while the average transaction value declined following the interstate banking deregulation, transaction values in sectors *more* dependent on external finance increased vis-à-vis transaction values in *less* external finance dependent sectors.

While we find that interstate banking deregulation has had an effect on the entry rate, the number of cross-border transactions, and the overall volume of multinationals' FDI, our analysis suggests that the intrastate bank branching deregulation had no significant impact on cross-border investment. These findings are consistent with Kerr and Nanda's (2009) work on the effects of the two banking deregulations on entrepreneurial activity and are suggestive of the importance of national banks versus single-state banks for FDI activity. Amore et al. (2013) find that interstate banking deregulation led to geographic diversification in the banking sector, which was beneficial for firms engaged in innovation.[7]

To study the impact of state-level banking deregulations on inbound FDI in the U.S. manufacturing sector, we employ transaction-level data collected by the International Trade

[7] Similarly, national banks may have a comparative advantage in evaluating foreign investment projects and multi-state banks may have better technology to serve multinational firms investing in the U.S. relative to single-state banks.

Administration (ITA) of the U.S. Department of Commerce. The ITA gathers data primarily from public sources, such as newspapers, trade journals, and public filings of federal regulatory agencies. The data identify the universe of new FDI transactions coming into the U.S. and contain information on the transaction value, the state where the foreign investment was made, the year of completion, and the nationality of the foreign investor.[8] The data also provide details on the type of transaction – e.g. new plant, merger and acquisition, or joint venture. We restrict our sample to transactions completed by 1994, which marks the passage of the 1994 Riegle-Neal Interstate Banking and Branching Efficiency Act that ended interstate banking and intrastate branching restrictions nationally.

We exploit time series variation in the adoption of intrastate branching and interstate banking deregulations across U.S. states to estimate the effect of cheaper local credit on the number and the size of new FDI transactions in the U.S. manufacturing sector. Formally, we specify a difference-in-differences econometric model with multiple time periods. Exploiting only *within* state variation in the two banking deregulations allows us to distinguish the effect of an increase in bank competition and the resulting reduction in the cost of borrowing from potential confounding factors. Because of the richness of the data, we are also able to control for a number of transaction- and investor-specific characteristics that may affect the average transaction value, such as the nationality of the foreign investor and the type of transaction. Our econometric models additionally include a host of state-level, time-varying covariates, such as the gross state product (and its growth rate), the unemployment rate, population density, the corporate tax rate, the average wage, the number of foreign trade zones, and market potential, all of which may affect FDI activity and be correlated with banking deregulation. Our results are robust to the inclusion of state-specific trends that additionally allow FDI trajectories to differ across states, as well as country-specific time effects and a host of variables characterizing investor experience. A major advantage of our study compared with cross-country studies is that we are implicitly able to control for

[8] In our data, the correlation between the ITA and the Bureau of Economic Analysis (BEA) measure of inbound FDI into the U.S. is 0.97. Similarly, Klein & Rosengren (1994) report a correlation between the two measures of 0.86 between 1979 and 1990.

many characteristics common to all states, such as macroeconomic policy and federal legislation (with respect to labor and capital markets as well as trade policy) that can affect FDI.

Our study contributes to a growing literature assessing the effects of credit constraints on international economic activity (e.g., Buch *et al.* 2009, 2010 and 2014, Manova 2008; Amiti & Weinstein 2011; Chor & Manova 2012).[9] The analysis presented here is most closely related to Klein *et al.* (2002) who find that changes in the supply of *source* country bank financing affects FDI activity for Japanese firms investing in the U.S. Our work complements theirs, as we show that access to *host* country external financing is just as important for the incidence and the intensity of FDI activity. Furthermore, our results are more comprehensive, as we use data on *all* FDI transactions into the U.S. manufacturing sector, regardless of the country of origin, and we provide evidence for the intensive as well as the extensive margin of FDI flows. Along similar lines to Klein *et al.* (2002), di Giovanni (2005) focuses on how the depth of the financial markets in the *source* country affects cross-border mergers and acquisitions (M&As). Employing cross-country data on M&As in the gravity equation econometric model, he finds that the size of financial markets has a strong positive association with domestic firms investing abroad. Antrás *et al.* (2009) develop a theoretical model of multinationals and imperfect capital markets, and demonstrate that weak financial institutions decrease the scale of multinational activity while simultaneously increasing the reliance on parent financing. In related work, Bilir *et al.* (2014) show that host country financial development affects the operation of U.S. multinationals investing abroad. Consistent with our findings, they document that more financially developed host countries attract more affiliates of U.S. multinationals and the effects are larger in magnitude for sectors that depend more heavily on external finance. The authors rationalize these patterns of U.S. multinationals investment abroad with a three-country model featuring financial frictions. By comparison, we take advantage of a

[9] Our work is also related to the literature on multinational firms that studies their global positioning strategies as a function of endowments, trade flows, and cross-border activity (see, for example, Markusen 1984; Blonigen 2005; Markusen & Venables 2007; Yeaple 2003a,b).

policy experiment, interstate banking and intrastate branching deregulation, to estimate the causal impact of cheaper local (host country) finance on the incidence and size of inbound FDI transactions across U.S. states.

The link between access to bank finance and real economic activity has been explored at length in the domestic context (e.g. Levine 2005). Cetorelli & Strahan (2006) and Kerr & Nanda (2009) have shown that firm entry and entrepreneurship among domestic firms react positively to banking deregulation. Michalski & Ors (2012) have shown that bilateral trade increased in state-pairs that liberalized their banking systems. Additionally, Cacciatore *et al.* (2015) study the domestic and international effects of banking deregulation in a two-country theoretical model that predicts increased domestic business entry, real appreciation, increased international borrowing and less pronounced business cycles. What is distinct about our study is that we focus on the effect of these same deregulations on foreign investment in the U.S. We find a similar effect on FDI activity – the entry rate and number of transactions increased and smaller value transactions became more prevalent, with the overall value of investment inflows being positive and growing.

The rest of the paper is structured as follows: Section 2 provides an overview of banking deregulations in the U.S. Sections 3 and 4 discuss the data and the econometric strategy, respectively. We present and discuss the results in Section 5. Section 6 concludes.

2. Banking Deregulation across U.S. States

Until the 1970s, banks in the U.S. were severely restricted by state statutes in their ability to expand across state borders and to branch within a state. Beginning in the late 1970s, states began allowing bank holding companies headquartered in other states, with which they had entered into reciprocal agreements, to acquire local banks (see Figure 1). The Riegle-Neal Interstate Banking and Branching Efficiency Act of 1994 deregulated interstate banking nationwide, except where individual states opted out, superseding

between-state agreements and effectively putting out-of-state banks on an equal footing with local banks (Kerr & Nanda 2009).[10]

Similarly, until the 1970s only a handful of states allowed unrestricted within state branching. Throughout the 1970s and 80s state branching law deregulation allowed banks to establish multiple branches within a state through mergers and acquisitions (M&As) and de novo branching. Since branching through M&As deregulation marks the leading edge of state branching deregulation reform (Cetorelli & Strahan 2006; Demyanyk *et al.* 2007), we use those dates to mark a state's adoption of intrastate branching deregulation.

Kroszner and Strahan (1999) argue that the timing of banking deregulation is related to the relative strength of private interest groups standing to gain from deregulation, e.g. large banks as well as small firms, which are dependent on bank finance. In addition to this private interest argument, Freeman (2002) and Berger *et al.* (2012) point out that the timing of banking deregulation is correlated with a state's past economic performance, while Huang (2008) suggests that the timing of deregulation could also be correlated with anticipated changes in future economic activity. It is unlikely that the timing of banking deregulation is directly linked to FDI lobbying, interests and economic activity. We check whether there is any systematic relationship between initial average FDI transaction value (as of 1977, the first year in our sample), as well as the FDI entry rate, and the year of deregulation. In unreported regression results we find that there is no economically or statistically significant relationship between initial FDI presence and the timing of the adoption of banking deregulations across states.

While many studies focus on intrastate branching deregulation alone (Jayaratne & Strahan 1996; Black & Strahan 2002; Berger *et al.* 2012), we explore the effect of *both* interstate banking and intrastate branching deregulation, similar to Black & Strahan (2002); Demyanyk *et al.* (2007), and Kerr & Nanda (2009). To study the effect of access to bank financing on inbound FDI, we exploit the staggered adoption of banking deregulation laws in the 48 contiguous states excluding Delaware and South Dakota, because

[10] Only Texas and Montana passed legislation to opt out of the interstate banking provisions of the Riegle-Neal Act before they were to go into effect in 1997 (Kroszner & Strahan 1999).

of the preponderance of credit card banks in these states (Black & Strahan 2002; Berger *et al.* 2012). The total value of commercial and industrial loans and the total value of FDI inflows into our sample states are depicted in Figure 2A. The increase in the total value of FDI inflows matches the expansion of total value of commercial and industrial loans in the 1980s. Moreover, the reduction and the recovery of FDI inflows during and after the 1990-1991 recession in the U.S. coincides with similar dynamics in the total value of loans.[11]

3. Data

To assess the impact of the two banking deregulations on the extensive and the intensive margin of inbound FDI, we use detailed, micro-level data on new inward foreign direct investment transactions in the U.S. manufacturing sector, across the 48 contiguous states, excluding Delaware and South Dakota, between 1977 and 1994. The starting point of our analysis is dictated by data availability, as FDI transaction data from the late 1960s and early 1970s are not available. The end point of our sample marks the passage of the 1994 Riegle-Neal Interstate Banking and Branching Efficiency Act – the federal regulation that ended state restrictions on bank expansions across local and interstate markets. Until 1994, the International Trade Administration (ITA) of the U.S. Department of Commerce was the federal agency that collected and disseminated micro-level data on FDI flowing into the U.S.[12] We manually collect the data from all annual print publications by the ITA. The ITA data cover the vast majority of inward FDI transactions that occurred in the U.S. (ITA 1977-1994). Information contained in the ITA data does not come from a mandatory survey but is primarily obtained from public sources, such as newspapers, magazines, trade journals, and public filings of federal regulatory agencies (e.g. the Securities and Exchange Commission, the Federal Trade Commission, and the Federal Reserve Board). The data include details on the transaction value, identity of the foreign investor (including country of

[11] The correlation between total FDI inflows and total commercial and industrial loans in Figure 2A is 0.61.
[12] As far as we are aware, after 1994, only the Bureau of Economic Analysis (BEA) collects such data, however their data are confidential and not publicly available.

origin), location of the investment (state) and the year the transaction was completed.[13] Each transaction is also classified into one of six modes of entry: merger and acquisition, new plant, plant expansion, equity increase, joint venture, and other. Panel B of Table 1 lists the percentage of transactions in each mode of entry category, as well as the top five investor countries. As Panel B shows, the vast majority of transactions involve financially developed source countries.[14]

To assess the effect of cheaper local credit on the extensive margin of FDI, we construct a state-level panel counting the number of new FDI transactions and FDI entry rates in each state-year cell. To analyze the impact of the two banking deregulations on the intensive margin of FDI, we employ the data on transaction values. To our knowledge, no prior research has analyzed the individual transaction-level data that include the transaction values. In related work, Klein *et al.* (2002) employ the ITA data on the subsample of FDI transactions originating from Japan between 1987 and 1994 to show that *source* country bank financing plays an important role for Japanese FDI projects in the U.S. Because the ITA sample of Japanese transactions with non-missing values is relatively small, Klein *et al.* (2002) focus on the number of transactions instead. Previous work on FDI has also employed a subsample of the state-level count data to analyze the U.S. location decision of foreign multinationals (Coughlin *et al.* 1991; Friedman *et al.* 1992; Friedman *et al.* 1996), or to assess the impact of environmental standards or labor regulations on FDI (Keller & Levinson 2002; Kandilov and Senses, forthcoming).

Importantly, the ITA data series on FDI are highly correlated with FDI data from the Bureau of Economic Analysis (BEA), which are based on confidential surveys and as such are considered more comprehensive. Figure 2B plots the data on total FDI inflows into the U.S. manufacturing sector from the

[13] The data provide information on the identity of the U.S. firm involved in the transaction if, for example, the transaction was a merger and acquisition or a joint venture. The location is most commonly listed as the state where the investment occurred, however, some transactions provide more detailed location coordinates, such as the city/town or county.

[14] While all of top five investor countries are considered financially developed (see e.g., Demirguc Kunt and Levine (1999)), there is considerable variation in the investors' use and reliance on credit in these countries. For example, domestic credit given to private sector firms as a share of GDP for these top investor countries range between 68% in the U.K. (classified as a market-based financial system) and 160% in Japan (classified as a bank-based financial system). The ratios for Canada, Germany, and France are 82%, 83%, and 90%, respectively.

BEA, and the total volume of inward FDI measure from the ITA.[15] The correlation between the BEA measure of aggregate inward FDI and the ITA measure of total inward FDI in the manufacturing sector between 1980 and 1994 is 0.97.[16] While the publicly available BEA data do not contain information on FDI inflows at the state-level, they include data on the number of foreign multinational enterprises (MNEs), and the value of plant, property, and equipment (PPE) owned by the MNEs at the state-level. Panel C of Table 1 shows that the correlations between these state-level measures from the BEA and the number of transactions, as well as the total value of transactions, from the ITA are positive and range between 0.34 and 0.43.

About half of the transaction observations do not have a reported value, but there is no reason to believe that the data are not missing at random. Except for the transaction value, data on all other transaction characteristics are always recorded. We find little differences in the distribution of transaction covariates (such as location, year of completion, source country, and mode of entry) across the two groups of FDI projects – those with and those without reported transaction values. The pseudo-R^2 for a logistic regression with a dependent variable indicating if the observation has a reported transaction value and a set of independent variables that includes dummies for all transaction covariates (state, year of completion, source country, and mode of entry) is less than 0.10, indicating that there is likely little selection on these observables.

While there exist estimators that can use information from observations with a missing dependent variable, they are not implemented often in practice because the improvement is usually small. Therefore, in most cases researchers ignore observations with missing information (Wooldridge 2001). We proceed with analyzing the sample of transactions with recorded values, but we show, in two different ways, that the omission of transactions with missing values likely has little effect on the results. First, when we

[15] The BEA data on total FDI inflows in the U.S. manufacturing sector is calculated as the sum of the following series: FDI financial inflow transactions, FDI debt instruments inflows, FDI equity inflows, and reinvestment of earnings. These series are publicly available only at the aggregate level, starting in 1980. We calculate the ITA measure of total volume of inward FDI as the sum of the transaction values that are reported in each year.

[16] Similar to the correlation we find, Klein and Rosengren (2002) report that the correlation between the BEA measure of inward FDI and the ITA measure of inward FDI between 1979 and 1990 is 0.86.

analyze the extensive margin, we create two different transaction count datasets – one that counts all transactions in each state-year cell (and therefore is not affected by missing transaction value observations) and another that counts only transactions with recorded values. We then proceed to estimate the impact of the two banking deregulations on the entry rate of foreign multinationals and on the number of new FDI transactions using both of these datasets.[17] The estimated impacts of the banking deregulation reforms across the two datasets are very similar, suggesting that omitting transactions with missing values may not bias the estimates much. Second, assuming selection on observables, we use inverse probability weighting to demonstrate that the results along the intensive margin remain largely unchanged. These results are reported in Appendix Table A2.

The next section provides details of our econometric strategy and describes the different state-level time-varying covariates that may affect either the entry rate of foreign multinationals (as well as the number of new inbound FDI transactions) or the FDI transaction value. These include the gross state product (from the U.S. Bureau of Economic Analysis); the state unemployment rate (from the U.S. Bureau of Labor Statistics); the average wage (from the Current Population Survey, U.S. Census Bureau); the state corporate tax rate (from World Tax Database, Office of Tax Policy Research, University of Michigan); the number of foreign trade zones (from the U.S. Foreign-Trade Zones Board, International Trade Administration, U.S. Department of Commerce); a market potential variable, calculated for each state *s*, and year *t* as the sum of all (real) gross state products of other states *n*, in year *t*, discounted by their centroid distance from state *s* (i.e. $\text{Market Potential}_{st} = \sum_{n \neq s} \frac{\text{GSP}_{nt}}{\text{Distance}_{ns}}$); and population density calculated as state population divided by total land area.

Summary statistics for all variables included in our analysis are presented in Panels A and B of Table 1. On average, there are about 6.93 new FDI projects annually in the manufacturing sector (3.52 projects with recorded transaction values), corresponding to an entry rate of 0.0115 in the average state,

[17] Because data on the number of existing multinationals in the manufacturing sector (the denominator of the entry rate) at the state level over our sample period between 1977 and 1994 are not available, we use data from the BEA on the number of all multinationals in manufacturing as well as non-manufacturing sectors in the state. About one half of employment in foreign owned firms in the U.S. is in the manufacturing sector.

also with significant variation across states – the minimum number of transactions is 0 and the maximum 103. The average transaction value over the sample period is $70.07 million (1983 dollars), but there is considerable variation – the smallest transaction is only $67,500 while the largest is over $7 billion.

Further, in Table 2, we present state-level data on the changes in U.S. commercial bank lending to foreign firms (changes in the 3-year average before and after the interstate banking deregulation) along with data on the changes in foreign investment activity (number of deals and average deal value) and data on structural changes in the state's banking sector (average loan yield, deposits HHI, and large bank assets ratio).[18] To construct a number of the variables in Table 2, we use additional data from the Federal Deposit Insurance Corporation (FDIC) and (bank-level) data from the Federal Financial Institutions Examinations Council's Consolidated Reports of Condition and Income (Call Reports). In the first two columns of Table 2, we also report the total value and the number of foreign transactions over the entire sample period (1977-1994) so as to highlight the relative importance of each state. The data in Table 2 show that there is a positive correlation (0.27) between the change in foreign loans and the number of FDI transactions as well as a negative correlation between the average loan yield and the number of transactions. This suggests that cheaper credit and larger loans to foreign firms may lead to more frequent foreign investments. We formalize this link in our econometric strategy below.

4. Econometric Strategy

4.1 Econometric Model for the Entry Rate of Foreign Multinationals and the Number of FDI Transactions

To assess the effect of the two banking deregulations on the extensive margin of FDI, we consider both the entry rate of foreign multinationals and the number of new inbound FDI transactions. We define the

[18] Loans to foreign addressees may under-report loans to U.S. affiliates of foreign multinationals since these affiliates have existing U.S. addresses rather than foreign addresses. Therefore, our analysis puts emphasis on the variables capturing structural changes in the banking industry: the average loan yield, deposits HHI, and the large bank assets ratio.

entry rate as the number of new inbound FDI transactions normalized by the total number of multinationals present in each state. Note that the entry rate we construct is the *gross* rate, as it does not take into account firm exits because exits are not reflected in our data. While typically the entry rate is positive, about twenty percent of all state-year observations of the number of new FDI transactions, and hence the entry rate of multinationals, are zeros. To accommodate for this, we specify a Tobit model, which is typically used both for censored regression applications and corner solution models. In the first instance, the dependent variable is censored above or below a certain value, for example as a result of the survey design. In the second instance, which is the case here, the dependent variable is a choice made by an agent. The dependent variable may take on a value of zero with positive probability because the optimal choice by the agent is a corner solution at zero but it is a continuous random variable over strictly positive values. In either case, it may be problematic to use Ordinary Least Squares (see Wooldridge 2001). Formally, we estimate the following Tobit model:

$$(1) \qquad \text{Entry\_Rate}^{*}_{st} = \alpha_1 \text{Interstate\_Bank}_{st} + \alpha_2 \text{Intrastate\_Branch}_{st} + \mathbf{X_{st}}\mu + \omega_s + \tau_t + \omega_s * Trend_t + \kappa_{st}$$

$$(2) \qquad \text{Entry\_Rate}_{st} = \max\{0, \text{Entry\_Rate}^{*}_{st}\},$$

where $\text{Entry\_Rate}^{*}_{st}$ is the underlying latent variable, which is not observed, and it satisfies the classical linear model assumptions and Entry_Rate_{st} is the observed outcome, defined as the number of new inbound FDI transactions in state *s* and in year *t* divided by the total number of foreign multinationals operating in that state and year (see the Data section and footnote 17). Equations (1) and (2) above imply that the observed variable, Entry_Rate_{st}, equals $\text{Entry\_Rate}^{*}_{st}$ when $\text{Entry\_Rate}^{*}_{st} \geq 0$, and $\text{Entry\_Rate}_{st} = 0$ when $\text{Entry\_Rate}^{*}_{st} < 0$.

The two indicator variables $\text{Intrastate\_Branching}_{st}$ and $\text{Interstate\_Banking}_{st}$ in equation (1) equal to one starting in the year in which each respective state allowed statewide bank branching and interstate banking, respectively, and zero otherwise. Our econometric model also includes a host of time-varying,

state-specific control variables that are likely to affect incoming FDI and may be correlated with banking deregulation. These controls are collected in the vector $\mathbf{X_{st}}$ and include three proxies for market size (demand) – (1) the natural logarithm of the gross state product for state *s* in year *t*, (2) the natural logarithm of the state's market potential (calculated for each state *s* in year *t*, as the sum of all neighboring states' real gross state products at time *t*, discounted by their centroid distance from state *s*, see the Data section for more details), and (3) population density for state *s* in year *t*; three proxies for the local cost of doing business – (4) the natural logarithm of the average wage, (5) the state corporate tax rate and (6) the number of foreign trade zones (FTZs) in state *s* in year *t*; and finally, (7) the unemployment rate in state *s* in year *t*, and (8) the current and lagged values of the growth rate of gross state product, which describe local economic conditions, and may be correlated with the timing of the adoption of banking deregulation (Freeman 2002; Huang 2008; Berger *et al.* 2012).[19]

In addition to the control variables listed above, we include state fixed effects ω_s, in order to control for unobservable, time-invariant, state-specific characteristics that affect the entry rate and the number of FDI transactions and may be correlated with the two bank branching deregulations. We also include year fixed effects τ_t, to capture economy-wide shocks that affect all states. Finally, to allow for cross-state differences in trends of FDI flows, we also include state-specific time trends $\omega_s * Trend_t$. It is important to account for such differences in trends since productivity growth differs across states, which could affect the investment decisions of foreign investors. Moreover, differences in productivity growth across states may be correlated with the adoption of the intrastate branching and interstate banking deregulations (Freeman 2002; Berger *et al.* 2012).

[19] Huang (2008) implements an alternative estimator that relies on the geographic discontinuity of intrastate banking deregulations. He compares the economic performance of contiguous counties that are separated by a state border, where intrastate branching restrictions exist only on one side of the border. This type of "geographic matching" is not suitable for our context for at least two reasons. First, the majority of FDI transactions lack county-level geographic information (almost all transactions do have information on the state). Second, most FDI transactions involve enterprises that are not located in counties along the state border, which rarely contain major cities or centers of economic development. Hence, such analysis cannot generalize to the entire state and its economy (Berger *et al.* 2012).

We estimate the Tobit model using maximum likelihood. The standard errors are adjusted for heteroskedasticity and are clustered by state. We weight all of the empirical specifications by the natural logarithm of the average state manufacturing employment in foreign multinationals over the period 1977-1985 (see, for example, Kerr and Nanda 2009).[20] Note that these weights are time-invariant and hence are not affected by the two banking deregulations over time. The weights are used in order to produce population estimates of the treatment effects of banking deregulation. We obtain economically and statistically similar results in unweighted regressions.

In addition to estimating the impact of the two banking deregulations on the entry rate of foreign multinationals, we also evaluate their effect on the number of new inbound FDI transactions. For this purpose, we specify a zero-inflated negative binomial model (see the Technical Appendix for details, also see Wooldridge 2001), which is a commonly used count data model with several advantages over the basic Poisson model or the computationally simpler negative binomial model that is not zero-inflated. We opt for a negative binomial instead of a Poisson model, in order to circumvent the mean-variance assumption of the latter (Cameron & Trivedi 1998). We fit a zero-inflated count model to avoid bias resulting from the large number of state-year cells with zero inbound FDI transactions. Note that while analyzing the number of new transactions in conjunction with the entry rate is informative, we focus most of our attention on the entry rate because it accounts for the existing presence of foreign multinationals when evaluating the effect of the two banking deregulations on FDI activity. Hence, looking at the effect of the two banking deregulations on the entry rate may be more meaningful since the same absolute change in the number of new inbound FDI transactions may be economically more important in states with smaller numbers of existing foreign firms.

[20] Data on state manufacturing employment in foreign multinationals are available from the BEA. Note that data on the number of foreign multinationals operating in the manufacturing sector are not available at the state level (see footnote 17).

4.2 Econometric Model for the FDI Transaction Values

To investigate the impact of the two banking deregulations on the value of FDI transactions in the U.S. manufacturing sector, we specify the following differences-in-differences econometric model with multiple time periods:

$$(3) \quad \ln V_{imcstj} = \beta_1 \text{Interstate\_Bank}_{st} + \beta_2 \text{Intrastate\_Branch}_{st} + \mathbf{X_{st}}\gamma + \mathbf{Z}_i\alpha + \omega_s + \tau_t + \delta_m + \\ + \lambda_c + \psi_j + \omega_s * Trend_t + \varepsilon_{imcstj},$$

where $\ln V_{imcstj}$ is the natural logarithm of the value (expressed in 1983 U.S. dollars) of transaction *i*, in mode of entry *m*, from source country *c*, in state *s*, in year *t*, and in two-digit SIC industry *j*. The vector $\mathbf{X_{st}}$ contains the state-specific, time-varying controls described in the previous subsection.

The vector $\mathbf{Z}_i$ includes four investor and transaction-specific covariates. First, we allow transaction values to systematically differ for investors that have invested multiple times in the U.S. Specifically, we include an indicator variable that takes on a value of one for investors that have completed multiple FDI transactions in the U.S. during our sample period, and zero for single-transaction investors. Multiple-transaction investors can be larger companies that run large scale operations, leading them to invest in higher-value projects. Second, having made prior investments in the U.S. can affect subsequent transaction values. On the one hand, a higher number of previous transactions would imply greater exposure to the local market, potentially increasing the value of subsequent transactions. To account for this market exposure effect, we additionally include a variable that counts the number of previous transactions. On the other hand, having invested previously implies that the foreign firm has already paid the sunk cost of entering the U.S. market, which could lower the average value of subsequent transactions. To capture this effect, we include a dummy variable that takes on a value of one if the foreign firm has previously invested in the U.S. and zero otherwise. Finally, we also include a variable that equals the natural logarithm of the average value of all previous investments and equals zero if this is the first transaction for the investor or there are no reported values for previous transactions. A higher

average value for previous investment transactions may indicate a high-value investor, so one would expect higher past averages to translate to higher current transaction values. However, higher past transaction value averages may also signal that the investor has already completed most necessary high-value investments, such as building a new plant or acquiring a large stake in a domestic company, and all that remains to be done are smaller adjustments, such as modest plant expansions or an incremental change in the ownership stake in the local company. In this case, current investment transactions will have lower values than the average of previous transactions for the investor.

In addition to the control variables listed above, our econometric model features a number of fixed effects. First, as in the model for the entry rate and the number of new transactions, we include state fixed effects, ω_s, year fixed effects, τ_t, and state-specific time trends, $\omega_s * Trend_t$. Mode of entry fixed effects, δ_m, are added to control for possible correlation between the value of the FDI transactions and the type of investment the foreign firm undertakes. For instance, the average value of a merger and acquisition transaction (106 million 1983 U.S. $) is similar to the average value of an equity increase transaction, but is about three times as large as a new plant transaction (about 28 million 1983 U.S. $) and about twice as large as a joint venture transaction (45 million 1983 U.S. dollars). Further, source country fixed effects, λ_c, are included to capture time-invariant, country-specific characteristics, such as the geographic distance from source country *c* to state *s* as well as legal and linguistic differences between the source country and the U.S. that affect the size of the FDI transaction. Because the value of firms across industries within the manufacturing sector may differ as a result of variation in productivity or market structure, we also include two-digit SIC industry fixed effects, ψ_j, to capture potential cross-industry differences in the value of new FDI transactions. Finally, the last term in our regression equation (3), ε_{imcstj}, denotes the residual.

Bertrand *et al.* (2004) show that inferences in a difference-in-differences setup with multiple time periods that combines micro-level data with state-level variation in regulations can be problematic due to

serial correlation issues. To address such concerns, we follow their suggestion and use heteroskedasticity robust standard errors that are clustered by state. This estimator of the variance-covariance matrix is consistent in the presence of any correlation pattern within states over time. As we do in the case for the extensive margin, we weight the empirical specification by the natural logarithm of the average state manufacturing employment in foreign multinationals over the period 1977-1985. Qualitatively and quantitatively similar results are obtained in unweighted regressions.

Finally, before presenting the results from our main specification above, we provide a visual summary of the data on entry rates and the natural logarithm of the average transaction value in Figures 3A and 3B.[21] These figures represent the estimates of a dynamic difference-in-differences model employing a set of dummies that measure the distance in years from the deregulation passage. The omitted category (reference group) is the year prior to deregulation. While we do not expect that there are any pre-treatment trends in the two outcomes of interest (the entry rate and the average transaction value), the two figures allow us to check that this is indeed the case. They also allow us to see if there are any delayed effects from the deregulations. Figure 3A provides preliminary evidence that following the interstate banking deregulation, the entry rate rose significantly. There are no discernable effects following the intrastate branching deregulation. Figure 3B suggests that the average transaction value

[21] Specifically, the figures plot the coefficients and the 95 percent confidence bands from estimates of the parameters α and λ from the following equation:

$$
\begin{aligned}
Outcome_{st} = {} & \alpha_1 \text{Interstate\_Bank}_{s,\{-10^+,-9,-8\}} + \alpha_2 \text{Interstate\_Bank}_{s,\{-7,-6\}} + \alpha_3 \text{Interstate\_Bank}_{s,\{-5,-4\}} + \alpha_4 \text{Interstate\_Bank}_{s,\{-3,-2\}} \\
& + \alpha_5 \text{Interstate\_Bank}_{s,\{0,1\}} + \alpha_6 \text{Interstate\_Bank}_{s,\{2,3\}} + \alpha_7 \text{Interstate\_Bank}_{s,\{4,5,6\}} + \alpha_8 \text{Interstate\_Bank}_{s,\{7,8,9,10^+\}} + \\
& + \lambda_1 \text{Intrastate\_Branch}_{s,\{-10^+,-9,-8\}} + \lambda_2 \text{Intrastate\_Branch}_{s,\{-7,-6\}} + \lambda_3 \text{Intrastate\_Branch}_{s,\{-5,-4\}} + \lambda_4 \text{Intrastate\_Branch}_{s,\{-3,-2\}} \\
& + \lambda_5 \text{Intrastate\_Branch}_{s,\{0,1\}} + \lambda_6 \text{Intrastate\_Branch}_{s,\{2,3\}} + \lambda_7 \text{Intrastate\_Branch}_{s,\{4,5,6\}} + \lambda_8 \text{Intrastate\_Branch}_{s,\{7,8,9,10^+\}} + \\
& + \mathbf{X_{st}}\gamma + \omega_s + \tau_t + \nu_{st},
\end{aligned}
$$

where $Outcome_{st}$ is the outcome of interest (the entry rate or the logarithm of the average transaction value) in state *s* in year *t*, and ω_s and τ_t are state and year effects, respectively; $\text{Interstate\_Bank}_{s,\{t\}}$ and $\text{Intrastate\_Branch}_{s,\{t\}}$ are dummy variables that assume the value of one only in years {t} relative to the adoption year, e.g. $\text{Interstate\_Bank}_{s,\{-5,-4\}}$ is a dummy variable equal to one five or four years prior to state *s* adoption of interstate banking. As before, $\mathbf{X_{st}}$ is a vector of state-level covariates. Huber-White standard errors allow for arbitrary error correlations within states.

declined after the passage of the interstate banking deregulation, while intrastate branching again had no noticeable effect.

5. Results

This section presents our empirical results. We first consider the extensive margin, and describe our estimates of the impact of the two banking deregulations on the entry rate of foreign multinationals, and the number of new inbound FDI transactions. We then turn to the impact of the two banking deregulations on the average value of inbound FDI transactions. Subsequently, we analyze how the transaction values change with previous investment history, and how banking deregulations affect the total value of foreign investments undertaken by multinationals that invest multiple times in the U.S.

5.1 Impact on the Entry Rate of Foreign Multinationals and on the Number of New Inbound FDI Transactions

In Table 3, we start by presenting the results from the Tobit specification (see equations (1) and (2)) that estimates the impact of the interstate banking and intrastate branching deregulations on the entry rate of foreign multinationals, i.e. on the extensive margin. In the first four columns we consider the entry rate computed using all new transactions (those with and those without recorded transaction values) in the numerator, and in column (5) we consider only the number of new transactions with recorded transaction values. In both cases, the denominator is the total number of existing multinationals operating in that state and year. Columns (1) - (3) and (5) report the results from the Tobit model, whereas column (4) reports the estimates from an equivalent OLS model for comparison. We report the average marginal effects for the Tobit specifications, so that the coefficients are directly comparable to the OLS estimates. Column (6) of Table 3 presents the results from the zero-inflated negative binomial model (see the Technical Appendix) that estimates the impact of interstate banking and intrastate branching deregulations on the *number* of new FDI transactions. We only report results using the total number of

new transactions (those with and without information on transaction values) to conserve space; the estimates using only transactions with non-missing values are both economically and statistically similar to the ones reported in Table 3.

Column (1) presents the results from the baseline Tobit specification for the entry rate without covariates. Using this specification, we obtain a positive and highly significant coefficient of 0.48 (with a standard error of 0.18) on interstate banking, which implies that the entry rate of foreign multinationals increased by 0.48 percentage points following the adoption of the interstate banking deregulation. Since the average entry rate (including transactions with missing transaction values) is 1.15 percent (see Table 1, Panel A), a 0.48 percentage point increase implies that, on average, states experienced an increase in the entry rate of about 42 percent (= 0.48/1.15) annually after interstate banking deregulation occurred. This estimate suggests that cheaper credit as a result of the interstate banking deregulation leads to a higher incidence of FDI transactions. Hence, *host* country financing is important for foreign multinationals. This finding complements previous work on FDI and exports by Klein *et al.* (2002) and Amiti and Weinstein (2011), who show that *home* country financing is important for Japanese FDI projects and firm exports, respectively.

Unlike the coefficient on the interstate banking indicator, the coefficient on intrastate branching is positive but insignificant.[22] The lack of a significant effect from the intrastate bank branching deregulation is aligned with the findings of Kerr and Nanda (2009), who document that while interstate banking brought about significant growth in entrepreneurship, as well as business closures across states, intrastate branching had little effect. The authors hypothesize that the result could be due to intrastate branching having a smaller impact on competition in the banking sector, or to multi-state banks having

[22] The correlation between the interstate banking and the intrastate branching indicators in our sample is 0.5. To ensure that the insignificance of the intrastate branching indicator is not driven by a potential multicollinearity problem, we re-estimated our empirical specification with the deregulation indicators included one at a time. Compared to the results reported in column (1) of Table 3, there is almost no change in the estimated coefficients on intrastate branching and interstate banking when the indicator variables are included one at a time. The coefficient on intrastate branching remains positive and insignificant, while the coefficient on interstate banking remains positive and significant.

the technology to serve new start-ups better than single-state banks. The latter argument also applies to multinational companies investing abroad. Furthermore, national banks may have a comparative advantage relative to single-state banks in evaluating foreign investment projects. The main channel through which interstate banking deregulation affects the incidence of inbound FDI, however, is most likely the availability of cheaper credit following the deregulation. We provide direct evidence for this channel later in this section.

When we augment the baseline specification with additional covariates and state trends, the coefficient on interstate banking does not change much and remains both statistically and economically significant. The estimate on interstate banking in our preferred specification in column (3) suggests that states that deregulated interstate banking experienced a 0.38 percentage point (or about 33 percent) increase in the FDI entry rate. There is little change in the estimated effect of interstate banking on the entry rate if we use OLS instead of the Tobit model. The OLS counterpart of the Tobit specification in column (3) of Table 3 is presented in column (4). The impact estimated with OLS is 0.45 (with a standard error of 0.18).

The results are similar if we instead consider the entry rate measure that uses only the number of transactions for which a transaction value is recorded. The average marginal effects of interstate banking in column (5) is quite similar to those in columns (1) - (4), but the estimated percentage increase is larger since the average entry rates computed when we use only transactions for which the value is not missing are about twice as low. In general, the banking deregulation impact on the entry rate of foreign multinationals reported in Table 3 is larger than what previous studies have documented for the extensive margin of domestic firms. For example, Black and Strahan (2002) find that the number of new incorporations per capita increased by 11 percent following interstate banking deregulation, and Kerr and Nanda (2009) find that the number of new single-unit start-ups and multi-unit facility expansions increased by 6 percent and 3 percent, respectively. One reason for the larger interstate banking deregulation impact on the entry rate of foreign multinationals is that their average entry rate of 1.15

percent (see Table 1) is considerably smaller than the entry rate of new domestic firms, which averages about 6 percent per annum (Lee *et al.* 2012). Therefore, even a small absolute change in the number of new FDI projects constitutes a large percentage change in terms of the entry rate. Another potential explanation for this difference is the footloose nature of multinational companies compared to their domestic counterparts (Caves 1996).

In column (6), we consider the total number of new inbound FDI transactions. Consistent with our findings from the specifications for the entry rate in the previous columns, the coefficient on the interstate banking indicator is both statistically and economically significant, but the one on intrastate branching is not. The estimated coefficient of 0.17 (with a standard error of 0.09) on interstate banking implies that the number of new inbound FDI transactions increased by 19 percent following the adoption of the interstate banking deregulation.[23] Since the average number of new transactions (including those with missing transaction values) is 6.93 (see Table 1, Panel A), a 19 percent increase implies that the average state recorded 1.28 additional transactions annually after interstate banking deregulation was adopted.

Next, we analyze some of the potential mechanisms through which banking deregulation can affect investment by foreign multinationals. To this end, we consider three changes in the banking industry brought about by banking deregulation (see, for example, Jayaratne and Strahan (1998)). First, we ask whether banking deregulation increased the incidence of FDI by reducing the cost of credit to foreign multinationals. As a measure of the cost of credit, we construct the average loan yield in state *s* during year *t*, by dividing all banks' total interest income on loans and leases by their total loans and leases given out in that state and year. We obtain these data from the Federal Deposit Insurance Corporation (FDIC). Second, we also check if enhanced bank competition as a result of banking deregulation affects foreign direct investment. To measure competition, we calculate the Herfindahl-

[23] Because the indicator variable only changes discontinuously, the effect of the interstate banking deregulation is calculated as $(e^{0.17}-1) = 0.19$. For estimated coefficients that are small in magnitude, this procedure makes little difference.

Hirschmann index (HHI) of bank deposit concentration using bank-level data from the Federal Financial Institutions Examinations Council's Consolidated Reports of Condition and Income (Call Reports).[24] Lastly, we consider the changes in the size distribution of banks following deregulation. As noted before, by allowing banks to acquire branches in different states, the interstate banking deregulation led to larger national banks that may have a comparative advantage in evaluating foreign investment projects. In order to capture the changes in the bank size distribution, we calculate the share of assets held by large banks in a given state and year using data from the Call Reports referenced above. We classify banks as "large" if their assets (measured in 1983 real dollars) in a given year are higher than the 75th percentile of the bank asset distribution for the full sample.

Panel A of Table 4 presents the effects of the interstate banking and the intrastate branching deregulations on the three banking industry measures discussed above. The estimates in column (1) show that the average loan yields declined significantly following the interstate banking deregulation. The estimated coefficient on interstate banking suggests that the average cost of loans declined by 7 percent. While the coefficient on intrastate branching is also negative, it is not statistically significant. The results for bank concentration in column (2) suggest that both interstate banking and intrastate branching significantly enhanced the competition in the banking industry, as bank deposit concentration decreased.[25] At the same time, the estimates in column (3) suggest that both deregulations led to an increase in the share of assets held by large banks, which is a result of bank consolidation following the deregulations.[26]

In Panel B of Table 4, we investigate how these structural changes in the banking industry affected the entry rate of foreign multinationals. Because inbound FDI and local banking industry characteristics such as loan yields and bank competition are likely endogenously determined, we take advantage of the policy experiment represented by the interstate banking deregulation to devise a standard

[24] We follow Jayaratne and Strahan (1998) in constructing the bank concentration measure, and calculate the HHI as the sum of the squared shares of statewide bank deposits held by each bank.

[25] The HHI is bounded between 0 and 1, with 0 implying perfect competition and 1 implying a monopoly. Hence, a reduction in HHI suggest an increase in the degree of competition in the industry.

[26] Bank competition increased, even though the share of assets held by large banks grew, as a greater number of large banks were allowed to compete in regional and state markets.

two-stage Instrumental Variables (IV) strategy, where we utilize the exogenous variation in the interstate banking deregulation and use it as an instrument for the changes in the local banking environment (loan yields, concentration, and bank size distribution).[27] We have already reported the first-stage of this IV strategy in Panel A of Table 4, where the estimates confirmed that all three of the banking variables we considered were affected by the interstate banking deregulation.[28] Panel B of Table 4 reports the two-stage IV estimates (where both equations are estimated simultaneously). Columns (1) and (2) report the results for the logarithm of the average loan yield. The coefficient on average loan yields is negative and statistically significant at 5 percent in the baseline specification without covariates, and it is significant at the 10 percent when we include additional state covariates in column (2). Focusing on the specification in column (2), the estimate shows that a one percent reduction in the average loan yield brought about by banking deregulation, led to a 0.11 percentage point increase in the entry rate of multinationals. Given the estimate of interstate banking deregulation in Panel A, our results suggest that cheaper credit provided an incentive for multinationals to invest in states that deregulated their banking system, and increased the entry rate of FDI by about 0.77 percentage points (= (-7 percent)*(-0.11)).

In columns (3) and (4) of Panel B in Table 4, we consider the impact of changes in banking industry competition, measured by the Herfindahl-Hirschmann index (HHI) of bank deposit concentration, on the entry rate of foreign multinationals. Similar to the specifications in the previous columns, we instrument for the log of HHI with the interstate banking deregulation indicator. As expected, the coefficient on the banking concentration variable is negative, suggesting that the entry rate of multinationals increased following the improvement in the banking competition (a reduction in the HHI) brought about by the deregulation. Specifically, we find that banking deregulation led to a 2 percent increase in banking industry competition (see Panel A of Table 4, column (2)), which in turn boosted the entry rate by about 0.56 percentage points (= (-2 percent)*(-0.28)) .

[27] We obtain very similar results if we use both interstate banking and intrastate branching indicators as instruments. The results are available upon request.

[28] F-test of the exclusion restriction in each of the first-stage IV regressions showed that the interstate banking regulation is an important determinant of changes in the loan yield, banking concentration and the bank size distribution (in all three cases, the F statistics was larger than 10).

In the last two columns of Panel B, we examine the role of changes in the bank size distribution. The coefficient on the large banks' assets ratio is positive and significant at the 1 percent level in the baseline specification, and significant at the 10 percent level in the broader specification with state-level time-varying covariates. These findings lend support to the hypothesis that by allowing bank consolidations, the interstate banking deregulation led to the formation of larger national banks that likely have comparative advantage in evaluating foreign investment projects, which, in turn, led to greater number of foreign transactions. Quantitatively, these estimates, combined with the impact of the interstate banking deregulation on the large banks' assets ratio in Panel A (column (3)), suggest that the banking deregulation increased the entry rate of foreign multinationals by about 0.88 percentage points (= 2 percent*0.44).

To provide further evidence that cheaper credit is an important channel through which interstate banking deregulation affects foreign direct investment in the U.S., we next show that manufacturing industries which are more reliant on external finance (Rajan & Zingales 1998) experienced a larger increase in the incidence of FDI following the deregulation. To this end, we categorize all FDI transactions into two groups – those in industries that are *more* dependent on external finance and those in industries that are *less* dependent on external finance – based on a measure of external finance dependence as defined in Cetorelli and Strahan (1998).[29] The external finance dependence variable takes on a negative value when the median firm in a two-digit SIC industry has free cash flow, and therefore is less external finance dependent, and a positive value when the median firm in an industry must issue debt or equity to finance investment. We construct an external finance dependence dummy variable that takes on a value of one when the transaction belongs to a more external finance dependent industry, and zero otherwise. We formally test whether the impact of banking deregulation varies with the degree of external finance dependence by including interaction terms between the external finance dependence dummy and

[29] Cetorelli and Strahan (2006) calculate the external finance dependence variable for each two-digit SIC industry as the median value of the proportion of capital expenditures financed with external funds, using data for Compustat firms over the 1980-1997 period.

the interstate banking and intrastate branching indicators in our preferred Tobit specification, which includes covariates and state trends, for the entry rates.[30]

In column (1) of Panel A of Table 8, the coefficient on the interstate banking and intrastate branching indicators capture the effect of deregulation on the entry rate in the *less* external finance dependent industries, and the interaction terms capture the additional effect of the deregulation on the *more* external finance dependent industries. The coefficient on interstate banking is marginally positive, albeit not significant, while its interaction with the external finance dependence dummy is positive and highly significant. These results provide further direct evidence of the importance of the local finance channel for foreign investment in the U.S. The estimates suggest that the entry rate of multinationals increased in all industries, but that the increase was far more pronounced in the *more* external finance dependent industries. Hence, by providing cheaper credit from larger national banks, the deregulation allowed a larger number of foreign multinationals that rely on external finance to invest in the U.S. As in all of our previous specifications, the coefficient on intrastate branching is very small and not significant. Its interaction with the external finance dependence indicator is positive, but also not significant. As expected, the effect is larger for *more* external finance dependent industries since facilitating access to local finance stands to benefit those sectors more.[31]

In column (2) of Table 8, we additionally check the robustness of our results to the inclusion of the real effective exchange rate. Our sample (1977-1994) covers the well-documented episode of the large U.S. dollar depreciation (see e.g. Froot and Stein (1991) and Blonigen (1997)). Depreciation of the dollar during the early 1980's could have stimulated foreign investment, perhaps even differentially so in

[30] Note that we cannot interact the deregulation measures with the continuous measure of external finance dependence in this specification, since we are calculating two entry rates for each state-year cell—one for *less* external finance dependent industries and one for *more* external finance dependent industries. Also, we allow the coefficients on the covariates to be different for the less and more external finance dependent industries by interacting them with the external finance dependence dummy variable.

[31] We also used two alternative measures of financial vulnerability (see, e.g. Manova, 2008) – tangibility and capital intensity – in place of external finance dependence. These estimates are in agreement with the conclusions we draw from the external finance dependence results presented above. Because none of the interaction terms with these two alternative measures are individually significant at the 5 percent level, these estimates are not reported to conserve space but they are available from the authors upon request.

external finance dependent industries, by making U.S. assets cheaper when evaluated in the currency of the foreign investor.[32] If the timing of the banking deregulation is correlated with the exchange rate depreciation, our estimates of the importance of local credit conditions may be over-stated. Since we exploit the within state variation in identifying the effects of banking deregulation and the real exchange rate is common to all states, we have already implicitly controlled for it in the baseline state-level entry rate specification. However, we can include an interaction term between the real effective exchange rate, calculated as the value of the U.S. dollar against a weighted average of foreign currencies, adjusted by a price deflator, and the external finance dependence dummy in the extended specification in Table 8.[33] The results in column (2) of Table 8 show that even when we control for the real effective exchange rate (via its interaction with the external finance dependence dummy), the coefficient on the interstate banking deregulation and its interaction with the external finance dependence dummy remain very similar to the estimates in column (1) and remain significant at the 5 percent level. Thus, it is unlikely that the positive effect of the interstate banking deregulation on the entry rate we uncover, especially for the external finance dependent industries, is explained by the depreciation episode in our sample.

Another potential channel through which interstate banking (but not necessarily intrastate branching) may affect the incidence of FDI transactions is the emerging economies of scale from the opportunity to exploit a relationship with one large national bank following the deregulation. In this case, foreign multinationals that have transactions with branches of the same bank in multiple states may experience larger impacts. To check this, we re-estimate our baseline specification for the entry rate by additionally including interaction terms between the deregulation (interstate banking and intrastate branching) and an indicator for multi-state investors. The results, which we report in column (1) of Panel B in Table 8 show that the impact of the interstate banking deregulation is indeed more pronounced for multinationals that invest across multiple states and as conjectured, there appear to be no differential

[32] Blonigen (1997) uses data from Japanese acquisitions in the U.S. to show that real dollar depreciation increases the incidence of such acquisitions, especially in industries with more firm-specific assets.
[33] The source for the real effective exchange rate data is the International Financial Statistics, IMF.

effects on the entry rate from intrastate branching. Because the estimates are not as precise (the interaction term between interstate banking and multi-state investor is statistically significant at the 10 percent level), we take these results as suggestive but not conclusive.

5.2 Impact on the Average FDI Transaction Value

Table 5 reports results from estimating the impact of the interstate banking and intrastate branching deregulations on the natural logarithm of the value of foreign direct investment transactions. Column (1) of Table 5 presents the results from our most basic specification that includes only the deregulation indicators along with a full set of state and year fixed effects. Using this specification, we obtain a negative and highly significant coefficient of -0.35 (with a standard error of 0.12) on the interstate banking deregulation indicator, suggesting that the average FDI transaction value declined by 29.5 percent following he adoption of the interstate banking deregulation.[34] This finding suggests that cheaper credit extended by larger national banks as a result of the interstate banking deregulation allowed smaller FDI transactions to take place because of the lower cost of obtaining credit. This, in turn, lowered the average value of foreign direct investment transactions in the U.S. Our evidence is consistent with prior work by Cetorelli and Strahan (2006), who find that higher bank competition following bank branching deregulation increased the share of small firms in the U.S. manufacturing industry. As in the case for the extensive margin, the coefficient on the intrastate branching indicator is not economically or statistically significant. As we did with the entry rate, we show below suggestive evidence consistent with the idea that cheaper local credit and economies of scale (lower cost of obtaining credit) encouraged FDI by examining the investment patterns of multiple transaction investors doing business across multiple states.

In the second column of Table 5, we present the results of the augmented specification that includes a set of time-varying, state-specific determinants of foreign investment transactions, which may be correlated with the adoption of the two banking deregulations. The estimated coefficient of -0.32 (with

[34] Because the dependent variable is expressed in logarithmic form and the indicator variable only changes discontinuously, the effect of the interstate banking deregulation is calculated as $(e^{-0.35}-1) = -0.295$.

a standard error of 0.11) on interstate banking deregulation remains very similar to the baseline specification without covariates in column (1), and is still significant at the 1 percent level. The estimate implies that the average transaction value declined by 27.4 percent in states that adopted the interstate banking deregulation.

Turning to the covariates included in our augmented specification in column (2), we find two of the variables that proxy for local costs – the natural logarithm of the wage rate and the state corporate tax rate— to be negative and statistically significant at the 1 percent level. Moreover, as expected, we find the number of foreign trade zones in a state (which provide incentives to foreign commerce) to be positive and significant at the 5 percent level. These results are not surprising as multinational businesses often consider local labor costs, tax laws and incentives as important factors in their foreign investment decisions. As suggested by Berger *et al.* (2012) and Freeman (2002), the current and lagged value of the growth rate of gross state product (GSP) control for the possibility that the two banking deregulations are correlated with current or past economic performance. While both the current and lagged values of the growth rate are positive as expected, they are not significant. The coefficient on the unemployment rate, which also captures the economic performance of the states, is marginally positive and not significant. Among the covariates that control for market size and agglomeration, the natural logarithm of current GSP is positive as expected, albeit statistically insignificant. On the other hand, contrary to what one would expect, the natural logarithm of market potential is negative and the population density has a very small negative coefficient, however, neither is statistically significant.

In columns (3)-(6) of Table 5, we progressively expand our empirical model to include state-specific time trends, source country, mode of entry and two-digit SIC industry fixed effects. The coefficient on the interstate banking indicator remains significant at the 5 percent level in all of the specifications, and its value declines slightly as more fixed effects are included. In column (7), we further expand our specification to include investor- and transaction-specific covariates, in addition to the state-specific covariates, the full set of fixed effects, and the state trends. The coefficient on interstate banking

increases in magnitude to -0.29 (with a standard deviation of 0.12), and is significant at the 5 percent level, suggesting that the passage of the interstate banking deregulation lead to a 25.2 percent decline in the average FDI transaction value. The multiple-transaction investor dummy variable is positive and significant, indicating that transactions completed by investors who undertake multiple projects are on average about 86.9 (= ($e^{0.62}$-1)*100) percent larger. The other variables pertaining to previous investment behavior are all positive, with the average value of previous investments being statistically significant.

In column (8), we additionally control for the logarithm of the bilateral real exchange rate between the source country's currency and the U.S. dollar.[35] The coefficients on the banking deregulation coefficients remain identical to the estimates in column (7). While not being significant, the positive coefficient on the real exchange rate suggests that an appreciation of the dollar (an increase in the exchange rate) will lead to an increase in the average transaction value, as the U.S. assets will become more expensive. Lastly, we include source country-specific year effects for the top five foreign investor countries and a composite category that includes all other countries, in addition to the full set of fixed effects (see Table 1 for the list of countries).[36] These investor country-by-year effects control for an exhaustive set of changes in foreign economic conditions, such as interest rates or exchange rates, and policies in the source country, such as changes in tax rates that can influence foreign firms' investment decisions. Focusing on this augmented specification in column (9), we find the coefficient on interstate banking deregulation to be -0.25 (with a standard error of 0.09). This estimate suggests that by increasing the availability of cheaper credit, interstate banking deregulation allowed smaller transactions to be financed and led to a 22.1 percent decline in the average FDI transaction value.

As we discussed in the data section, about half of the transaction observations do not have a reported value. While there is no reason to believe that the data on transaction values are not missing at

[35] The bilateral real exchange rate is constructed as the official nominal exchange between the foreign currency and the U.S. dollar, adjusted by the CPI's of the source country and the U.S. (source: International Monetary Fund's International Financial Statistics (IFS)). The exchange rate and the CPI data are not available for a few of the countries in our sample. We use the U.S. real effective exchange rate for those countries.

[36] FDI inflows from the top five countries make up 80 percent of the total number of transactions in our sample. Each of the other countries in our sample has very few transactions, which makes it difficult to identify those country-specific year effects.

random, especially conditional on all fixed effects, state trends, and covariates, we provide additional evidence that this does not affect the estimated coefficients much. Assuming selection on observables, we use inverse probability weighting to demonstrate that the results along the intensive margin remain largely unchanged. To estimate the weights, i.e. the propensity score of having a recorded transaction value and therefore being included in the analysis, we specify a logistic regression with a dependent variable indicating if the observation has a reported transaction value and a set of independent variables that includes dummies for all transaction covariates (state, year of deal completion, source country, and mode of entry). We then use the inverse of the predicted probability (the propensity score) as weights in regression equation (3). The corresponding results are presented in Appendix Table A2. The estimated effects of both banking deregulations are very similar to those reported in Table 5, where we do not correct for missing values.

In Table 6, we perform a number of additional robustness checks. Because of the detailed nature of our data, when we analyze individual transaction values, we can also include investor-specific fixed effects. Note that when we do include investor fixed effects, identification of the coefficients on the two banking deregulations only uses within-investor variation in deregulation, i.e. only data on multiple-transaction investors are employed. Columns (1) and (2) of Table 6 present the results for the single and multiple-transaction investors separately.[37] While the coefficient on interstate banking for the single-transaction investors subsample is negative and very similar in magnitude to the one obtained for the full-sample, the coefficient estimated with the multiple-transaction investor subsample is larger in magnitude and highly statistically significant (-0.43 with a standard error of 0.13). The estimated impact remains unchanged when we include investor-specific fixed effects, suggesting that the average transaction value for foreign firms investing multiple times in the U.S. declined by 35.6 (= $(e^{-0.44}-1)*100$) percent following interstate banking deregulation.[38]

[37] We estimate the augmented specification (3), which corresponds to column (7) of Table 5.

[38] We also find that both single- and multiple-transaction investors experienced an increase in the number of completed transactions and entry rates, but the effect is significantly larger in magnitude for the multiple-transaction investors. These results are available upon request.

These results (column (3) of Table 6) further suggest that for a given multiple-transaction investor, transaction size declined with the number of previous investments, indicating that, on average, investors undertake major projects upon entering the U.S. market and subsequently make smaller adjustments. The results are consistent with the idea that interstate banking deregulation created economies of scale for multiple-transaction investors as they were enabled to consolidate their banking relationships across states. This evidence dovetails with anecdotal reports suggesting that as a result of interstate banking deregulation, firms operating across multiple states could significantly reduce the number of individual bank relationships, potentially translating into major savings.[39]

In columns (4) and (5) of Table 6, we perform two additional robustness checks related to multiple-transaction investors by combining all transactions that were undertaken by the same foreign firm investing in the same target company in the same state and year into one aggregate transaction. There are few such instances of multiple transactions and combining those into one transaction produces results that are very similar to those using the original data – compare column (4) with column (2) of Table 6 and column (5) with column (3) of Table 6.

We next analyze how the changes in the banking industry's structure that emerged from banking deregulation impacted the average transaction value. As in the case for the entry rate specifications in Panel B of Table 5, we focus on the changes in the average loan yield, bank concentration, and the large banks' ratio of total assets in the state. As before, we use the interstate banking deregulation as an instrument in each case. The results in columns (1) and (2) of Table 7 show that there is a positive relationship between the cost of loans, measured by the average loan yield, and the average value of the foreign transactions. Focusing on the richer specification in column (2), the estimates imply that a one percent reduction in the average loan yield leads to a 4.67 percent decline in the average transaction value. Combining this effect with the impact of interstate banking deregulation on the average loan yield

[39] A CNN Moneyline report from the deregulation period suggests that the total savings in fees and staff time could amount to $400 million per year in 1994 dollars. As an example, the treasurer of a large publishing company, active in 41 states suggested that their firm could reduce their banking relationships from 300 to 50, leading to cost savings of $250,000 per year.

presented in Panel A of Table 4, our estimates show that interstate banking deregulation lowered the average transaction value by 32.69 percent by reducing the cost of loans. While the coefficient on the bank concentration measure is negative in columns (3) and (4), suggesting that higher concentration (less bank competition) is associated with lower transaction value, it is not significant. In the last two columns of Table 7, we consider the changes in the bank size distribution. We obtain negative coefficients on the large banks' assets ratio that are statistically significant at the 10 percent level both in the baseline and in the broader specification with state covariates. The estimate in column (6) suggests that a one percentage point increase in the large banks' assets ratio led to a 22.61 percent reduction in the average transaction value. Hence, by allowing for bank consolidations, interstate banking deregulation created larger banks that are likely in a better position to evaluate and fund smaller investment projects.

Finally, we test how the interstate banking and branching deregulations impact the average foreign transaction value in sectors more reliant on external finance differently than those in sectors that are less reliant on external finance. To do so, we include, in equation (3), interaction terms between the deregulation indicators and the continuous measure of external finance dependence of Cetorelli and Strahan (1998). Column (3) of Panel A in Table 8 presents these results. The main effect of interstate banking remains very similar to the findings in Table 5, and it is significant at the 1 percent level. The interaction term is positive and significant at the 5 percent level, which confirms that following banking deregulation, transaction values in *more* external finance dependent industries increased relative to transaction values in *less* external finance dependent industries as access to cheaper credit improved. As in all previous specifications, neither the interaction term nor the main effect of the intrastate branching deregulation is significant.

Another potential channel through which interstate banking (but not necessarily intrastate branching) may affect the incidence of FDI transactions is the economies of scale from the opportunity to exploit a relationship with one large national bank following the deregulation. Then, foreign multinationals that have transactions with branches of the same bank in multiple states may experience

larger impacts. As we did with the entry rate, to check this, we re-estimate our baseline specification for the transaction values by additionally including interaction terms between the deregulation (interstate banking and intrastate branching) and an indicator for multi-state investors. The estimates are reported in column (2) of Panel B in Table 8. They show that the effect of the interstate banking deregulation is indeed more pronounced for multinationals that invest across multiple states, but the coefficient is not precisely estimated.

5.3 Impact on Total Investment Volume of Foreign Multinationals

In this section, we evaluate the effect of the two banking deregulations on the total value of all investments in the U.S. undertaken by foreign multinationals. Our results for the extensive margin show that foreign firms increase the frequency of their FDI transactions following the adoption of the interstate banking deregulation. On the other hand, our results for the intensive margin show that as local access to finance improves as a result of interstate banking deregulation, foreign firms decrease the average transaction value. Hence, it is unclear if foreign multinationals increase the total value of their investment in the U.S.

To address this question, we analyze how the two banking deregulations affect foreign multinationals' overall investment in the U.S. To this end, we aggregate the transaction-level data up to the investor-level by adding up all transactions in a year completed by a given investor across different states, industries, and modes of entry. We then specify the following econometric model which estimates how the fraction of states that have deregulated interstate banking or intrastate branching affects foreign firms' overall investment in the U.S.:

$$(6)\ \ln V\_Total_{ict} = \beta_1 \text{Share\_Interstate\_Bank}_t + \beta_2 \text{Share\_Intrastate\_Branch}_t + \mathbf{Z}_{it}\alpha + \\ + GDP\_Growth_t + Time\_Trend_t + \lambda_c + \varepsilon_{ict}.$$

The dependent variable, $\ln V\_Total_{ict}$, is the natural logarithm of the overall investment (across all states, industries, and modes of entry) of foreign investor *i* from source country *c* in year *t*. The two variables Share_Interstate_Banking$_t$ and Share_Intrastate_Branching$_t$ represent the fraction of U.S. states that have deregulated interstate banking and intrastate bank branching, respectively. The vector $\boldsymbol{Z}_{it}$ includes investor-specific covariates, such as the number of previous investments. Our model also controls for the U.S. GDP growth (GDP_Growth_t) as FDI is typically pro-cyclical, a time trend ($Time\_Trend_t$) to account for any aggregate trend in foreign direct investment, as well as source country fixed effects (λ_c) to capture time-invariant, country-specific characteristics that affect foreign multinationals' overall investment in the U.S.

The coefficients of interest are β_1 and β_2, which estimate the impact of the share of states that have adopted the interstate banking deregulation or the intrastate branching deregulation on the overall U.S. investment undertaken by foreign multinationals. The results are presented in Table 9. Consistent with all of our previous results, we find that the interstate banking deregulation matters whereas the intrastate branching deregulation does not. In particular, we estimate that when the fraction of states which have adopted interstate banking deregulation increases, overall FDI investment rises as well. The impact is economically meaningful and statistically significant in all of our specifications from column (1) to column (6). Using the full sample of all investors with the most detailed specification but without investor fixed effects yields an estimate of 1.44, which implies that as the share of states that have adopted the interstate banking deregulation rises by 10 percent, which is equivalent to 5 additional states adopting the interstate banking deregulation, foreign multinationals increase their total investment in the U.S. by 14.4 percent. Given that the average real investment value per foreign investor in our sample is 82 million (1983 U.S. $) and that the average number of investors per year is 165, the 14.4 percent increase implies an increase in total FDI inflows into the U.S. manufacturing sector of 1.9 billion (1983 U.S. $). Since the annual average (over 1980-1994 sample) of total FDI inflows into the U.S. manufacturing sector is 23.2 billion (1983 U.S. $, reported by the BEA), a 10 percent increase in the share

of states that have adopted the interstate banking deregulation accounts for 8.2 percent of the total FDI inflows. Using the sample of multiple-transaction investors and including investor fixed effects in the model produces similar results – increasing the share of states that have adopted the interstate banking deregulation by 10 percent leads to 18.2 percent increase in foreign investment.

6. Conclusion

Following the 2008 global financial crisis, economists have been reminded yet again that access to finance is important to economic activity, both domestically and internationally. Our work contributes to the growing literature on the economic impact of access to cheaper credit in two important ways. First, we provide direct micro-level evidence from the U.S. that improved access to cheaper *local* credit affects both the *intensive* and the *extensive* margins of inbound FDI and significantly raises foreign multinationals' overall investment in the U.S. We do so by using transaction-level data on new FDI projects across U.S. states from 1977 until 1994 and analyzing the value of each transaction (the intensive margin) as well as the FDI entry rate (the extensive margin) in each state. Second, we extend the current literature by providing estimates of the impact of changes in the cost of *local* credit on foreign firms investing in the U.S. To do this, we employ cross-state variation in the timing of two financial deregulations, intrastate branching and interstate banking, that increased local banking competition and lowered the cost of local credit. We estimate a difference-in-differences model with multiple time periods and control for a number of investor- and transaction-specific covariates, such as the nationality of the foreign firm and the mode of entry (merger and acquisition, joint venture, etc.), as well as a host of state-specific, time-varying characteristics, including state trends, that can affect inbound FDI and can be correlated with the timing of the banking deregulations. An important advantage of our empirical setup, compared with a cross-country analysis, is that we are implicitly able to control for many characteristics common to all states, such as macroeconomic policy and federal legislation, with respect to labor and capital markets as well as trade policy, which can affect foreign direct investment.

Our results reveal that the interstate banking deregulation has had a significant impact on FDI inflows both along the extensive and the intensive margin – it is associated with a higher entry rate of foreign multinationals, a larger number of new inbound FDI projects, and a smaller average project value. Analyzing the impacts of the two banking deregulations along the extensive margin shows that following the interstate banking deregulation, states experienced an average of 19 percent increase in new FDI projects, which translates into an additional 1.28 new FDI transactions annually. Further, our estimates suggest that the entry rate of foreign multinationals in states, which deregulated interstate banking rose 0.38 percentage points from an average of 1.15 percent before to an average of 1.53 percent after the adoption of the interstate banking deregulation. Also, we estimate that the average value of foreign transactions declined by about 22.1 percent in states that adopted the interstate banking deregulation. This result, which is robust to including a comprehensive set of state-level, time-varying controls and trends, investor- and transaction-specific covariates, as well as source country and mode of entry fixed effects, suggests that when access to cheaper local credit improves, foreign firms are able to undertake smaller inbound FDI projects.

Taken together, these results imply that by facilitating access to cheaper local finance, the interstate banking deregulation has enabled more foreign firms to invest in the U.S. and allowed them to undertake smaller projects on average. Moreover, we also document an economically and statistically significant positive impact of the interstate banking deregulation on multinationals' overall investment volume in the U.S. – our estimates imply that as the share of U.S. states that allow interstate banking rose by 10 percent, foreign firms' overall investment in the U.S. increased by 14 to 18 percent. Finally, to shed light on the potential mechanisms behind the effect of the interstate banking deregulation, we show that lower cost of local credit and greater local bank competition in each state, resulting from the interstate banking deregulation, stimulated FDI activity. Moreover, we also demonstrate that the impact of the deregulation is greater for industries that are more heavily dependent on external finance. The evidence underscores the importance of local, host-country, credit markets for FDI flows.

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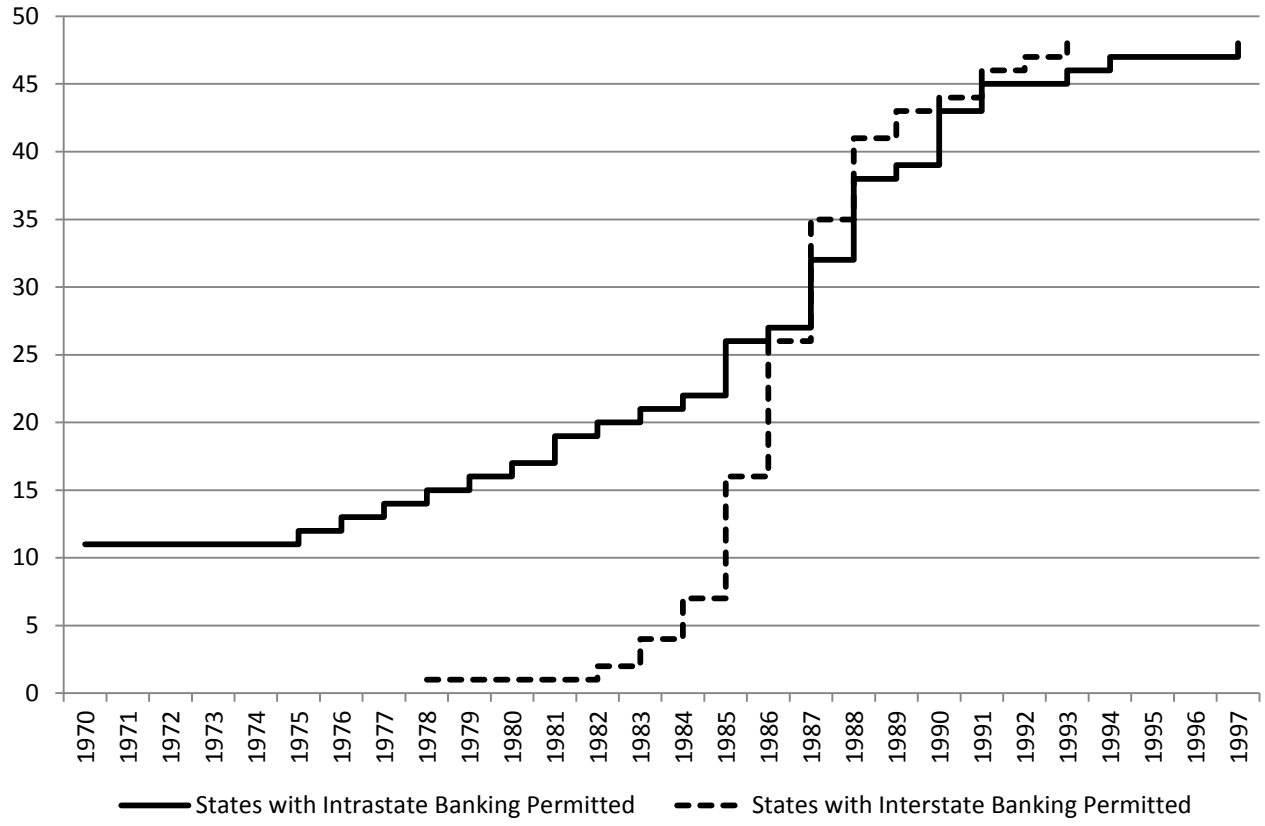


Figure 1. Number of states that deregulated intrastate banking and intratstae branching

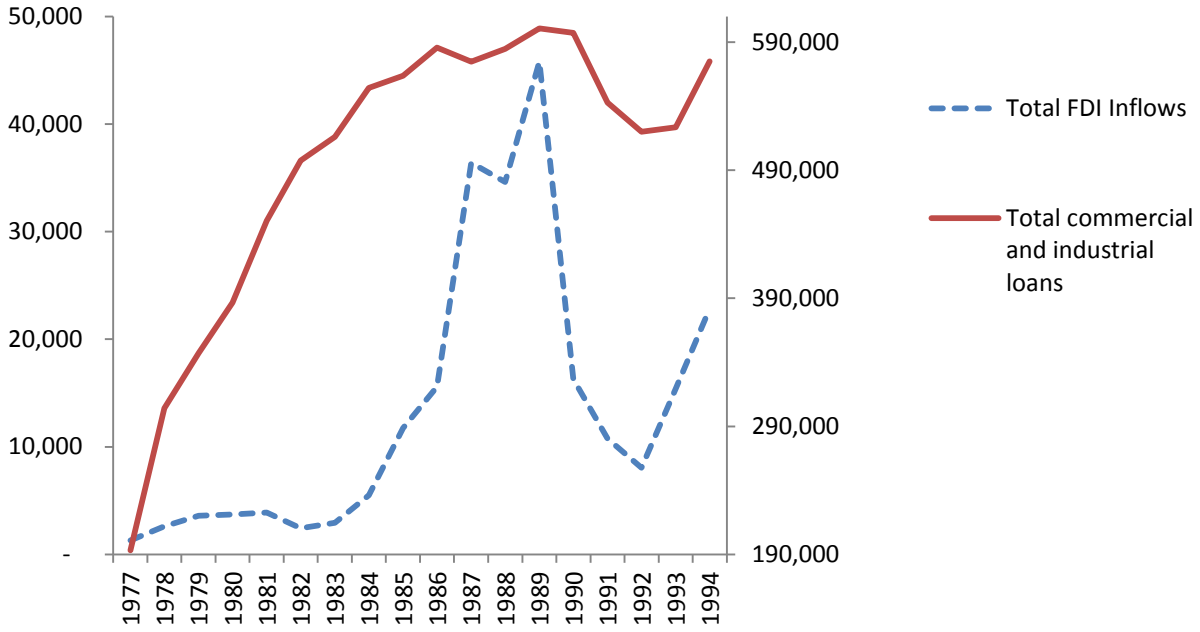


Figure 2A: FDI Inflows into the US Manufacturing Sector and Total Commercial and Industrial Loans. The figure plots the total value of the transactions for the 46 contiguous states in our sample and the total value of commercial and industrial loans given out by the banks in those states. The FDI data (left-axis) and the total loans data (right axis) are in millions of U.S. Dollars.

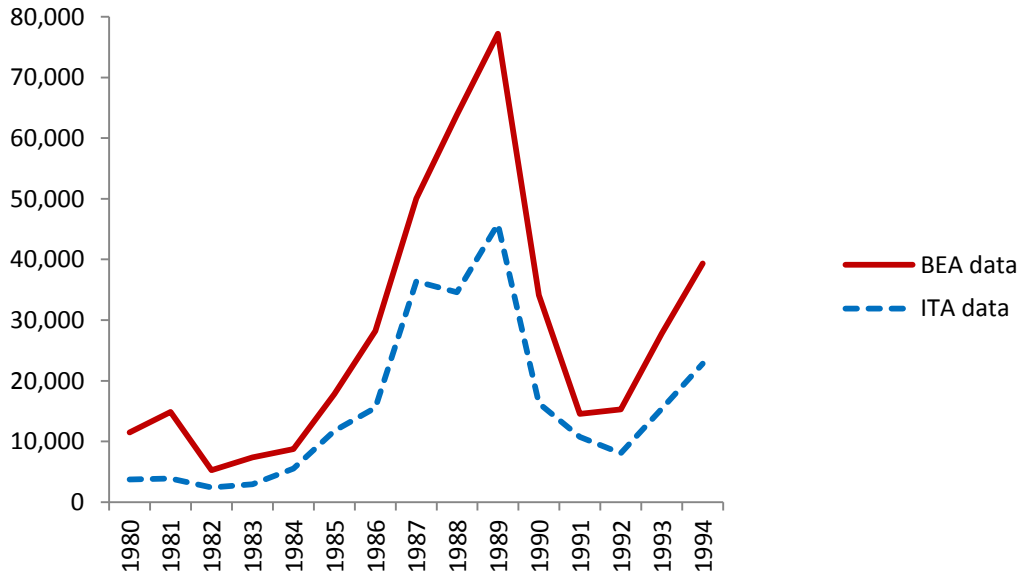
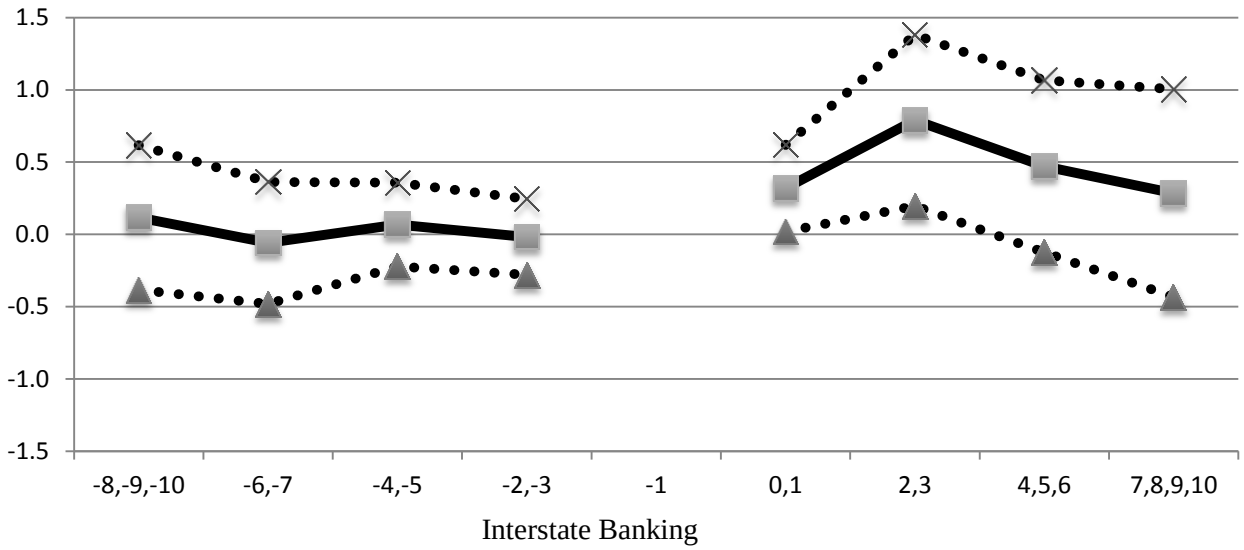
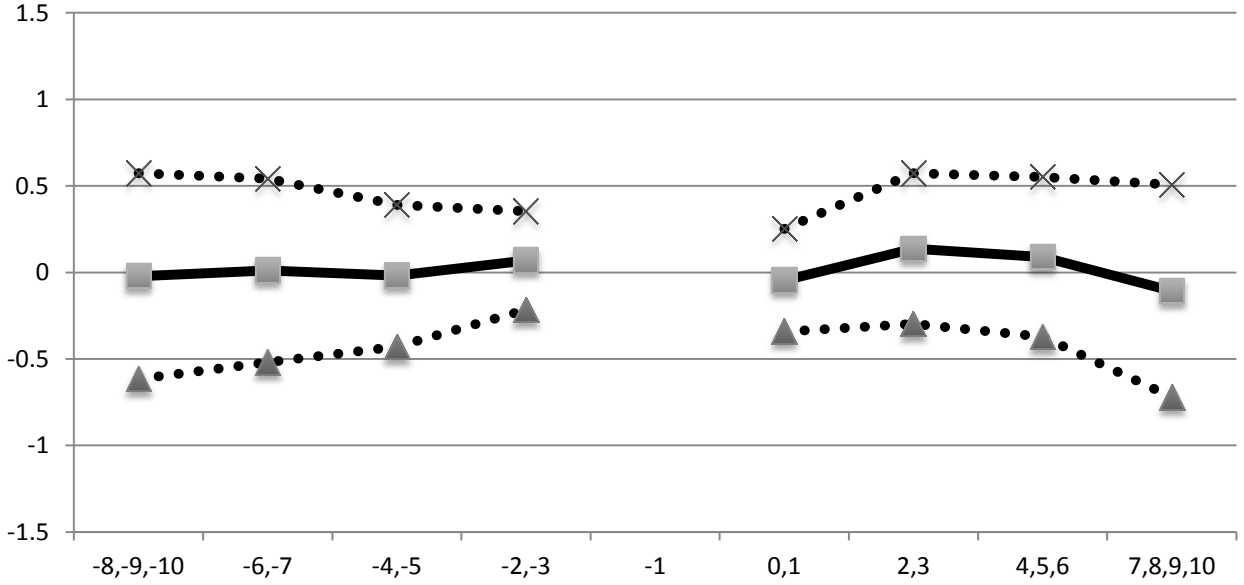


Figure 2B: FDI Inflows into the US Manufacturing Sector. The figure plots the total value of the transactions for the 46 contiguous states in our sample and the total value of FDI inflows into the U.S. manufacturing sector as reported by the Bureau of Economic Analysis (BEA). Both series are in millions of U.S. Dollars and the correlation between the two series is 0.97.

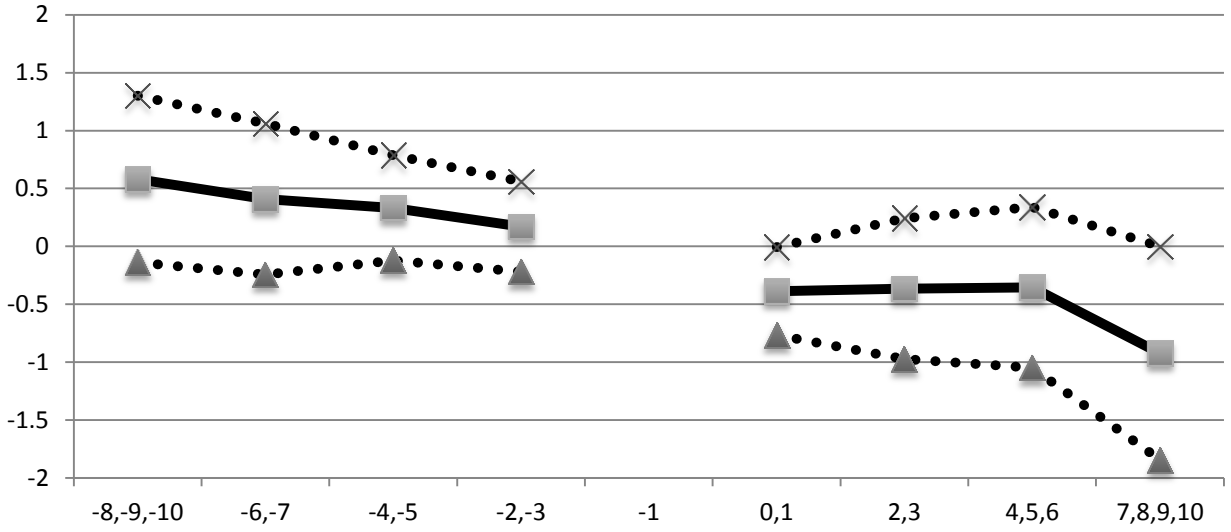


Interstate Banking

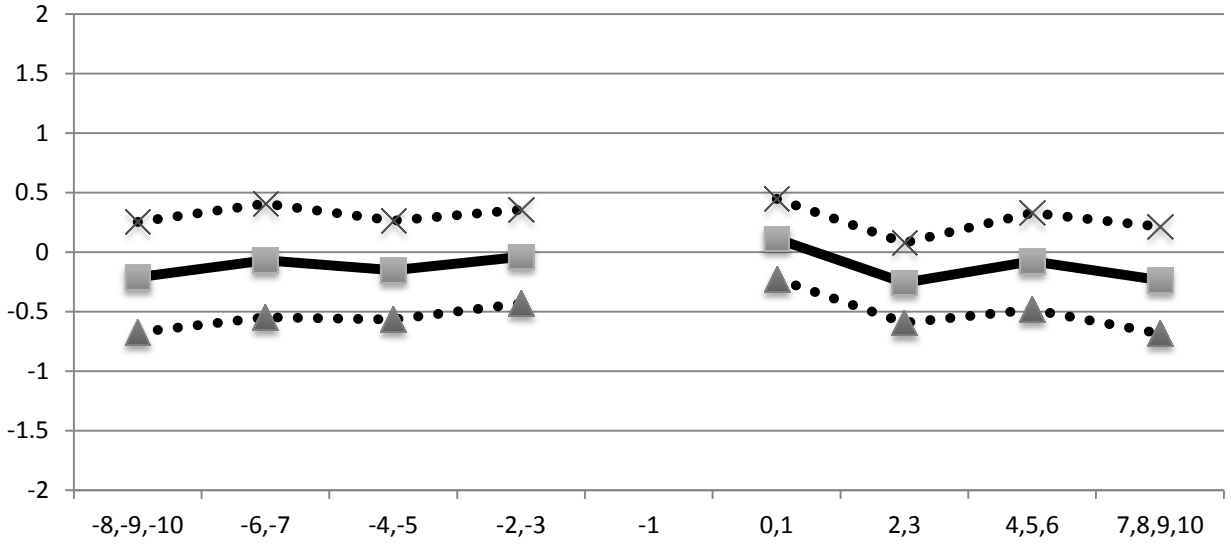


Intrastate Branching

Figure 3A. Intrastate and Interstate Banking Deregulation Dynamics for Entry Rates. The top and bottom panels track the dynamics for interstate banking and intrastate branching deregulation, respectively. The figures plot coefficients from regressions of log entry rates on a series of indicator variables extending from 10 years before the reform's passage to 10 years afterward. The end points include all earlier and later years. The indicator variable for the year before the reform is omitted, so that coefficients are measured relative to the year before the reform. Covariates, state and year fixed effects, as well as state trends are included in the regressions. The raw dynamics surrounding the passage of the intrastate and interstate reforms are estimated jointly. The dashed lines present 95% confidence intervals, with standard errors clustered by state.



Interstate Banking



Intrastate Branching

Figure 3B. Intrastate and Interstate Banking Deregulation Dynamics for Transaction Values. The top and bottom panels track the dynamics for interstate banking and intrastate branching deregulation, respectively. The figures plot coefficients from regressions of log values on a series of indicator variables extending from 10 years before the reform's passage to 10 years afterward. The end points include all earlier and later years. The indicator variable for the year before the reform is omitted, so that coefficients are measured relative to the year before the reform. Covariates, state and year fixed effects, as well as state trends are included in the regressions. The raw dynamics surrounding the passage of the intrastate and interstate reforms are estimated jointly. The dashed lines present 95% confidence intervals, with standard errors clustered by state.

Table 1. Summary Statistics. This table presents the summary statistics for the data used in our analysis. The inbound foreign direct investment transaction data come from annual publications of the International Trade Administration of the U.S. Department of Commerce. We use data on new inward foreign direct investment transactions in the U.S. manufacturing sector, across the 48 contiguous states, excluding Delaware and South Dakota, between 1977 and 1994. The total number of observations (for which the transaction value is not missing) is 2,915.

Panel A: Main Characteristics

Variable	Mean	St. Dev.	Min.	Median	Max
Transaction value (1983 U.S. $, millions)	70.07	269.88	0.07	12.51	7,035.21
Interstate banking	0.65	0.48	0.00	1.00	1.00
Intrastate branching	0.77	0.42	0.00	1.00	1.00
Average loan yield	0.15	0.03	0.09	0.15	0.26
Herfindahl-Hirschmann index of bank deposit concentration	0.04	0.07	0.00	0.02	0.33
Large banks' assets ratio	0.78	0.18	0.32	0.82	0.995
Gross state product (1983 U.S. $, millions)	196,998	161,719	5,523	146,598	591,783
Population density (persons per square mile)	255.69	246.28	7.29	169.73	1,080.86
Unemployment rate (percent)	6.58	1.75	2.28	6.32	15.57
Real wage (weekly, 1983 U.S. $)	861.28	612.52	20.05	723.02	2,890.49
State corporate tax (percent)	7.01	2.94	0.00	8.00	12.25
Number of foreign trade zones	4.28	3.93	0.00	3.00	27.00
Market potential (1983 U.S. $, millions)	584,235	224,114	159,625	622,923	1,292,423
Number of FDI transactions	6.93	10.46	0.00	3.00	103.00
Number of FDI transactions excluding those with missing values	3.52	5.47	0.00	1.00	58.00
Entry rate ×100	1.15	1.27	0.00	0.79	12.27
Entry rate excluding transactions with missing values ×100	0.60	0.75	0.00	0.40	5.45
Multiple-transaction investor indicator	0.50	0.50	0.00	1.00	1.00
Number of previous transactions for multiple transaction investors	1.34	3.10	0.00	0.00	22.00
Average previous transaction value for multiple transaction investors (1983 U.S. $, millions)	0.94	3.34	0.00	0.00	55.81

Table 1. Summary Statistics (Cont'd).

Panel B. Additional Characteristics

Variable	Percent of all transactions
Type of FDI Transactions	
- Mergers and acquisitions	46.8
- New plant	23.4
- Plant expansion	14.1
- Equity increase	7.7
- Joint venture	4.6
- Other	3.5
Nationality of foreign investor (top 5 countries)	
- Japan	36.6
- U.K.	19.2
- Germany	10.8
- Canada	9.3
- France	4.8
- Other	19.3

Panel C. Correlations: The table reports the correlation between the number of transactions and the total value of transactions (1983 U.S. $) for each state and year in our sample with the state-level data reported by the Bureau of Economic Analysis. The first row reports the correlations with the (1983 U.S. $) value of the plant property and equipment (PPE) owned by the foreign multinational enterprises (MNEs). The second row reports the correlations with the number of MNEs operating in a state. Number of state-year observations used in calculating the correlations is 748.

	Number of transactions	Total value of transactions
Total value of PPE	0.36	0.34
Number of MNEs	0.42	0.43

Table 2. Before-after deregulation comparison. Column (1) reports the total number of manufacturing transactions while column (2) reports the value of all manufacturing transactions with non-missing values in our sample. Columns (3)-(8) report three years before to three years after the deregulation changes in average key variables.

	Manufacturing FDI Transactions		Change (t-3, t+3)					
State	Number	Value (1983 U.S. $, millions)	Avg. No. New Deals	Avg. Trans. Value	Foreign Loans	Avg. Loan Yield	Deposits Herf.	Big Bank Assets Ratio
AL	63	15.71	0.58	-0.09	5.40	-0.020	0.021	0.063
AR	23	3.62	0.08	0.12	3.67	0.002	0.009	0.001
AZ	33	6.44	-0.50	0.17	-14.51	-0.020	0.000	-0.011
CA	833	197.97	20.33	-0.32	3,112.18	-0.013	0.033	0.001
CO	67	9.77	1.25	0.09	91.05	0.009	0.010	-0.027
CT	145	106.48	0.92	0.95	38.78	-0.030	-0.038	0.019
FL	138	39.27	3.75	0.57	360.73	-0.057	-0.006	0.055
GA	189	33.93	12.25	-0.26	-83.55	-0.046	-0.017	0.089
IA	35	6.49	-1.92	0.83	5.65	-0.040	0.012	0.037
ID	2	0.01	-0.33	0.00	-0.03	-0.027	-0.073	0.006
IL	260	126.49	13.58	0.71	690.14	-0.021	0.000	0.030
IN	146	30.11	15.67	0.29	42.18	-0.034	0.000	0.047
KS	32	4.16	0.67	0.09	0.05	-0.039	0.016	0.023
KY	122	52.58	9.67	0.46	0.51	-0.049	-0.017	0.042
LA	50	16.199	-1.58	0.16	-2.27	-0.013	0.014	0.012
MA	206	58.27	3.08	-0.24	-663.10	-0.071	-0.057	0.026
MD	84	16.90	4.33	0.09	-160.05	-0.046	-0.026	0.028
MI	231	70.64	16.50	-0.14	-396.47	-0.029	0.000	0.042
MN	66	68.35	3.83	1.58	-363.90	-0.015	0.000	0.021
MO	69	32.44	5.33	1.18	5.07	-0.035	0.000	0.080
MS	15	1.67	0.25	0.10	0.00	-0.005	0.036	0.023
NC	264	41.75	14.58	0.02	-70.64	-0.049	-0.047	0.019
ND	1	<0.01	-0.33	0.00	0.00	-0.044	0.027	0.064
NE	7	0.73	-0.42	0.00	-0.21	-0.025	0.027	0.063
NH	22	5.66	-0.83	0.40	-0.09	-0.005	0.010	0.091
NJ	300	96.16	5.33	0.28	29.86	-0.034	0.000	0.013
NM	11	0.86	0.42	0.00	-0.41	0.003	0.038	0.017
NV	58	42.38	10.50	0.86	0.00	-0.046	-0.112	0.009
NY	609	312.54	-2.75	-0.06	-13,133.80	-0.028	-0.050	0.000
OH	301	192.15	18.42	0.91	109.44	-0.050	-0.010	0.028
OK	29	12.82	1.67	1.20	3.30	0.002	0.003	-0.065
OR	65	13.70	7.25	-0.09	81.21	-0.020	0.000	0.015
PA	258	151.46	3.75	3.10	95.47	-0.029	0.000	0.012
RI	35	9.97	1.33	-0.11	73.98	-0.043	-0.132	0.005
SC	84	24.06	4.92	0.33	1.04	-0.040	0.000	0.027
TN	154	37.57	9.50	0.27	77.41	-0.051	-0.006	0.069
TX	319	130.14	1.42	0.32	227.70	0.006	0.012	0.008
UT	13	2.55	0.08	0.23	-2.34	-0.043	-0.104	0.016
VA	155	31.06	0.75	0.24	44.27	-0.039	-0.021	0.044
VT	17	0.31	0.08	0.01	0.33	0.003	0.035	0.067
WA	96	18.54	6.25	0.28	-290.77	-0.010	0.028	0.011
WI	88	28.89	-1.25	0.62	11.83	-0.021	0.004	0.062
WV	22	8.50	0.33	0.02	0.00	-0.015	0.008	0.071

Table 3. The Impact of Interstate Banking and Branching Deregulation on the Entry Rate of Foreign Multinationals and the Number of Transactions in the U.S. Manufacturing Sector, 1977-1994. The dependent variable in columns (1)-(5) is the FDI entry rate, defined as the fraction of new inbound FDI transactions normalized by the total number of multinationals present in each state. In columns (1)-(4) the entry rate is calculated using the number of new inbound FDI transactions with and without reported values. In column (5), the observations with missing values are excluded when calculating the entry rate. Columns (1)-(3) and (5) estimate Tobit models for the entry rate, and column (4) estimates an OLS model. Column (6) presents results from a zero-inflated negative binomial (ZINB) model for the number of new FDI transactions *N*, in each state *s* (including those with missing values). The following explanatory variables are normalized by 100: Interstate banking and Intrastate branching indicator variables, the unemployment rate, the corporate tax rate, the number of foreign trade zones, and the population density. All specifications include a full set of state and year fixed effects. Columns (3)-(6) additionally include state-specific time trends. All the specifications are weighted by the log average state manufacturing employment in foreign multinationals. Robust standard errors clustered at the state level are reported in parentheses. *** denotes significance at the 1 percent level, ** at the 5 percent level, and * at the 10 percent level.

Dependent Variable – Entry Rate (No. of New FDI Trans./No. of Multinationals)						No. of New FDI Trans
	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable:	Entry Rate (No. of New FDI Trans./No. of Multinationals)					No. of New FDI Trans
Interstate banking/100	0.48***	0.48***	0.38**	0.45**	0.37***	0.17**
	(0.18)	(0.18)	(0.18)	(0.18)	(0.14)	(0.09)
Intrastate branching/100	0.25	0.06	0.07	0.04	-0.04	0.03
	(0.16)	(0.14)	(0.19)	(0.19)	(0.17)	(0.11)
ln(GSP)		-0.03***	-0.01	0.01	0.00	-0.01
		(0.01)	(0.02)	(0.01)	(0.01)	(1.00)
GSP Growth rate		-0.03	-0.06***	-0.04**	-0.03**	-1.21
		(0.02)	(0.02)	(0.02)	(0.01)	(0.91)
GSP Growth rate lag		0.05**	0.03	0.02	0.02	-0.61
		(0.02)	(0.02)	(0.01)	(0.02)	(0.77)
Unemp rate/100		0.03	-0.07	-0.10**	0.02	-0.06
		(0.05)	(0.06)	(0.05)	(0.04)	(0.04)
ln(Wage)		0.04***	0.01	-0.01	0.01	-0.02
		(0.01)	(0.016)	(0.01)	(0.01)	(1.03)
Corporate tax/100		-0.03	-0.18	-0.19**	0.02	-0.00
		(0.09)	(0.13)	(0.10)	(0.10)	(0.04)
Foreign trade zone/100		-0.00	0.06	0.07	0.02	0.03
		(0.02)	(0.08)	(0.08)	(0.07)	(0.05)
ln(Market potential)		-0.04	-0.05	-0.11**	0.03	-0.03
		(0.06)	(0.07)	(0.05)	(0.05)	(2.78)
Population density/100		0.01	-0.03	-0.03	-0.04**	0.00
		(0.01)	(0.03)	(0.02)	(0.02)	(0.02)
State trends	No	No	Yes	Yes	Yes	Yes
Model specification	Tobit	Tobit	Tobit	OLS	Tobit	ZINB
Include missing value obs.	Yes	Yes	Yes	Yes	No	Yes
Log Likelihood/ R^2	19,636	19,962	20,442	0.60	19,663	-15,410
No. Obs.	828	828	828	828	828	828
Zero Obs.	183	183	183	183	262	183

Table 4.

Panel A: The Impact of Interstate Banking and Branching Deregulation on the Banking Industry. The table presents estimates of the effects of banking deregulation on various measures of banking industry's structure. The dependent variable in the first column is log of average loan yields, measured as the ratio of total interest income to the total loans given out by banks in a given state and year. In column (2), the dependent variable is the log of the Herfindahl index measuring bank concentration. In column (3), the dependent variable measures the fraction of total assets held by large. All specifications include year and state fixed effects, as well as state specific time-trends. All the specifications are weighted by the log average state manufacturing employment in foreign multinationals. Robust standard errors clustered at the state level are reported in parentheses. *** denotes significance at the 1 percent level, ** at the 5 percent level, and * at the 10 percent level.

	(1)	(2)	(3)
Dependent variable:	*ln*(Average loan yield)	*ln* (Bank concentration)	Large banks' assets ratio
Interstate banking	-0.07***	-0.02**	0.02***
	(0.02)	(0.01)	(0.01)
Intrastate branching	-0.02	-0.02**	0.02***
	(0.02)	(0.01)	(0.01)
R^2	0.92	0.76	0.99
No. Obs.	828	828	828

Panel B: The Impact of Changes in the Banking Industry's Structure on the Entry Rate of Foreign Multinationals in the U.S. Manufacturing Sector, 1977-1994. The table presents the estimates of the impact of three banking industry measures on the entry rate of foreign multinationals using Instrumental Variables estimation, where the measures are instrumented by the interstate banking deregulation measure. All specifications include year and state fixed effects, as well as state specific time-trends. All the specifications are weighted by the log average state manufacturing employment in foreign multinationals. Robust standard errors clustered at the state level are reported in parentheses. *** denotes significance at the 1 percent level, ** at the 5 percent level, and * at the 10 percent level.

	Dependent Variable-- Entry Rate (No. of New FDI Trans./No. of Multinationals)					
Variable	(1)	(2)	(3)	(4)	(5)	(6)
ln(Average loan yield)	-0.08**	-0.11*				
	(0.03)	(0.06)				
ln(Bank concentration)			-0.24**	-0.28		
			(0.11)	(0.18)		
Large banks' assets ratio					0.28***	0.44*
					(0.10)	(0.24)
Covariates	No	Yes	No	Yes	No	Yes
R^2	0.47	0.43	0.33	0.30	0.42	0.23
No. Obs.	828	828	828	828	828	828

Table 5. The Impact of Interstate Banking and Branching Deregulation on Average FDI Transaction Value in the U.S. Manufacturing Sector, 1977-1994. The dependent variable, $ln\ V_{imcst}$ is the natural logarithm of the real value of transaction *i*, in mode of entry *m*, from source country *c*, in state *s*, in year *t*, and in two-digit SIC industry *j*. The two indicator variables $Intrastate\_Branching_{st}$ and $Interstate\_Banking_{st}$ equal to one starting in the year in which the state allowed statewide bank branching and interstate banking, respectively, and zero otherwise. The specifications are weighted by the log of the average state manufacturing employment in foreign multinationals as weights. Column (1) presents results from a specification that includes the deregulation indicators along with a full set of state and year fixed effects. Column (2) presents the results of an augmented specification that includes a set of time-varying, state-specific determinants of foreign investment transactions. In columns (3)-(6) we progressively expand the empirical model to include state-specific time-trends, source country, mode of entry and two-digit SIC industry fixed effects. Column (7) reports results with investor- and transaction-specific covariates, in addition to the state-specific covariates, a full set of fixed effects. In column (8), we further include the log of the real exchange rate. In column (9), we replace the log of the exchange rate with source-country-specific year effects in addition to the full set of fixed effects. Robust standard errors clustered at the state level are reported in parentheses. *** denotes significance at the 1 percent level, ** at the 5 percent level, and * at the 10 percent level.

Dependent Variable – *ln*(Transaction value)

Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Interstate banking	-0.35***	-0.32***	-0.33**	-0.29**	-0.27**	-0.25**	-0.29**	-0.29**	-0.25***
	(0.12)	(0.11)	(0.13)	(0.14)	(0.13)	(0.13)	(0.12)	(0.12)	(0.09)
Intrastate branching	0.02	0.05	0.04	0.06	0.05	0.04	0.04	0.04	0.09
	(0.12)	(0.12)	(0.16)	(0.16)	(0.15)	(0.16)	(0.17)	(0.17)	(0.17)
ln(GSP)		0.93	0.58	0.68	1.07	0.90	1.05	1.07	1.74
		(0.93)	(1.70)	(1.61)	(1.52)	(1.39)	(1.56)	(1.55)	(1.64)
GSP Growth rate		1.25	1.04	0.96	0.72	0.59	-0.30	-0.33	-0.96
		(1.68)	(1.83)	(1.88)	(1.80)	(1.73)	(1.82)	(1.81)	(1.87)
GSP Growth rate lag		2.20	2.75	2.12	2.38	3.26	2.86	2.86	2.95
		(2.13)	(2.39)	(2.43)	(2.47)	(2.31)	(2.56)	(2.59)	(2.50)
Unemp rate		0.03	-0.02	-0.00	-0.01	-0.01	-0.01	-0.01	0.01
		(0.05)	(0.07)	(0.07)	(0.07)	(0.07)	(0.08)	(0.08)	(0.07)
ln(Wage)		-1.50***	-2.47	-2.09	-2.74	-2.49	-2.07	-2.06	-1.63
		(0.51)	(1.81)	(1.81)	(1.78)	(1.78)	(1.88)	(1.88)	(1.75)
Corporate tax		-0.13**	-0.16**	-0.15**	-0.17**	-0.18**	-0.16**	-0.16**	-0.20**
		(0.05)	(0.06)	(0.05)	(0.06)	(0.05)	(0.06)	(0.06)	(0.06)
Foreign trade zones		0.03**	-0.00	0.00	-0.01	0.01	-0.01	-0.01	-0.01
		(0.01)	(0.06)	(0.07)	(0.06)	(0.06)	(0.06)	(0.06)	(0.06)
ln(Market potential)		-0.70	-3.22	-2.28	-3.68	-2.43	-2.59	-2.61	-4.38
		(4.30)	(5.34)	(5.41)	(6.17)	(5.98)	(6.16)	(6.17)	(6.52)
Population density		-0.00	0.00	0.00	-0.00	-0.01	-0.01	-0.01	-0.02
		(0.00)	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)	(0.02)
Multiple-transaction investor							0.62***	0.62***	0.58***
							(0.08)	(0.08)	(0.08)
Previously invested							0.15	0.15	0.16
							(0.12)	(0.12)	(0.13)
No of previous investments							0.01	0.01	0.03**
							(0.01)	(0.01)	(0.01)
ln(avg. value of prev. invest.)							0.13***	0.13***	0.11***
							(0.03)	(0.03)	(0.02)
ln(real exchange rate)								0.03	
								(0.06)	
State trends	No	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Source country fixed effects	No	No	No	Yes	Yes	Yes	Yes	Yes	Yes
Mode of entry fixed effects	No	No	No	No	Yes	Yes	Yes	Yes	Yes
Two-digit industry fixed effects	No	No	No	No	No	Yes	Yes	Yes	Yes
Source country × year fixed effects	No	No	No	No	No	No	No	No	Yes
R^2	0.079	0.085	0.099	0.148	0.179	0.218	0.255	0.259	0.295
No. Obs.	2,915	2,915	2,915	2,915	2,915	2,915	2,915	2,915	2,915

Table 6. The Impact of Interstate Banking and Branching Deregulation on Average Foreign Direct Investment Transaction Value for Single versus Multiple Transaction Investors in the U.S. Manufacturing Sector, 1977-1994. We create two subsamples: one for single-transaction investors and the other for multiple-transaction investors, and estimate our augmented specification on these two subsamples. Columns (1) and (2) present the results for the single and multiple-transaction investors. Column (3) augments the Column (2) specification for multiple-transaction investors with investor-specific fixed effects. In columns (4) and (5), we check the robustness of the results for the multiple-transaction investors by aggregating transactions that were undertaken by the same foreign firm investing in the same target company in a given state and year. The results remain unchanged as a result of this aggregation. Robust standard errors clustered at the state level are reported in parentheses. *** denotes significance at the 1 percent level, ** at the 5 percent level, and * at the 10 percent level.

Dependent Variable – *ln*(Transaction value)

Variable	(1)	(2)	(3)	(4)	(5)
	Single-Transaction Investors	Multiple-Transaction Investors	Multiple-Transaction Investors	Multiple-Transaction Investors Aggregated Data	Multiple-Transaction Investors Aggregated Data
Interstate Banking	-0.27	-0.43***	-0.44**	-0.36**	-0.39**
	(0.18)	(0.13)	(0.19)	(0.15)	(0.19)
Intrastate Branching	0.02	0.02	-0.08	0.04	-0.07
	(0.24)	(0.28)	(0.26)	(0.28)	(0.25)
ln(GSP)	3.09	2.09	4.30*	1.45	4.41*
	(2.15)	(1.52)	(2.47)	(1.63)	(2.52)
GSP Growth rate	-2.42	1.71	-2.41	1.07	-3.09
	(3.21)	(2.99)	(4.53)	(2.94)	(4.77)
GSP Growth rate lag	7.39*	-2.15	-6.02	-2.16	-5.86
	(3.99)	(2.52)	(3.86)	(2.66)	(3.98)
Unemp rate	0.05	-0.06	-0.16*	-0.05	-0.16*
	(0.08)	(0.08)	(0.09)	(0.08)	(0.09)
ln(Wage)	-2.91	-3.13	-3.32	-2.19	-3.12
	(2.27)	(2.06)	(2.89)	(2.02)	(2.91)
Corporate tax	-0.16	-0.23**	-0.27*	-0.25**	-0.29*
	(0.10)	(0.10)	(0.16)	(0.10)	(0.16)
Foreign trade zones	-0.08	0.03	-0.00	0.07	0.00
	(0.06)	(0.09)	(0.08)	(0.09)	(0.08)
ln(Market potential)	-12.73	5.99	1.74	5.94	1.46
	(9.39)	(7.02)	(7.96)	(6.75)	(8.10)
Population density	0.01	-0.03	-0.07*	-0.02	-0.07**
	(0.04)	(0.03)	(0.03)	(0.03)	(0.03)
Previously invested		0.15	-0.88***	0.16	-0.85***
		(0.10)	(0.13)	(0.11)	(0.15)
Number of previous investments		0.03*	0.07***	0.02	0.07***
		(0.01)	(0.03)	(0.01)	(0.02)
ln(Average value of previous investments)		0.09***	-0.41***	0.10***	-0.39***
		(0.03)	(0.05)	(0.03)	(0.06)
Investor-specific fixed effects	No	No	Yes	No	Yes
R^2	0.32	0.35	0.64	0.35	0.63
No. Obs.	1,443	1,472	1,472	1,429	1,429

Table 7. The Impact of Changes in the Banking Industry's Structure on Average FDI Transaction Value in the U.S. Manufacturing Sector, 1977-1994. The table presents the results for the impact of the various banking industry measures on the average FDI transaction value using Instrumental Variables estimation, where the banking measures are instrumented by the interstate banking deregulation indicator. All specifications include year, state, source country, industry and transaction type fixed effects, as well as state specific time-trends and source country specific time effects. All the specifications are weighted by the log average state manufacturing employment in foreign multinationals. Robust standard errors clustered at the state level are reported in parentheses. *** denotes significance at the 1 percent level, ** at the 5 percent level, and * at the 10 percent level.

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)
ln(Average loan yield)	5.16**	4.67*				
	(2.11)	(2.33)				
ln(Bank concentration)			-33.42	-8.83		
			(56.04)	(6.72)		
Large banks' assets ratio					-32.50*	-22.61*
					(17.27)	(12.68)
ln(GSP)		2.82*		2.20		3.22
		(1.60)		(2.04)		(2.11)
GSP Growth rate		-3.55		-1.44		-0.39
		(2.48)		(2.44)		(2.13)
GSP Growth rate lag		3.60		3.85*		6.52*
		(2.35)		(2.05)		(3.59)
Unemp rate		-0.04		0.08		0.12
		(0.08)		(0.08)		(0.10)
ln(Wage)		-0.74		-3.54		1.25
		(2.01)		(2.55)		(3.09)
Corporate tax		-0.14*		-0.15*		-0.17**
		(0.08)		(0.08)		(0.07)
Foreign trade zone		0.01		0.01		0.00
		(0.07)		(0.06)		(0.06)
ln(Market potential)		-2.43		6.32		6.90
		(7.03)		(5.97)		(7.27)
Population density		-0.02		-0.03		-0.05
		(0.03)		(0.03)		(0.03)
Multiple-transaction investor		0.61***		0.56***		0.57***
		(0.08)		(0.08)		(0.08)
Previously invested		0.11		0.15		0.13
		(0.13)		(0.13)		(0.13)
Number of previous investments		0.03**		0.03**		0.03***
		(0.01)		(0.01)		(0.01)
ln(Avg. value of previous investments)		0.10***		0.10***		0.10***
		(0.02)		(0.02)		(0.03)
Observations	2,915	2,915	2,915	2,915	2,915	2,915
R-squared	0.07	0.29	0.00	0.28	0.00	0.27

Table 8. The Impact of Interstate Banking and Branching Deregulation on Average FDI Transaction Value and Entry Rate (No. of New FDI Trans./No. of Multinationals) – External Financial Dependence and Multi-state Investor Results; U.S. Manufacturing Sector, 1977-1994.

Panel A: We categorize all transactions into two groups—industries that are *more* external finance dependent and industries that are *less* external finance dependent— based on a measure of external finance dependence as defined in Cetorelli and Strahan (1998), and construct an external finance dummy that takes on a value of one when the transaction belongs to a more external finance dependent industry. Column (1) presents Tobit results for the entry rate of foreign multinationals (new FDI transactions/number of foreign multinationals) as dependent variable and includes interaction terms between the external finance dependence dummy and the interstate banking and intrastate branching indicators, normalizing the two indicator variables Interstate banking and Intrastate branching by 100. Note that we cannot interact the deregulation measures with the continuous measure of external finance dependence in this specification, since we are calculating two entry rates for each state-year cell—one for *less* external finance dependent industries and one for *more* external finance dependent industries. We allow the coefficients on the covariates to be different for the less and more external finance dependent industries by interacting them with the external finance dependence dummy variable. The specification includes state and year dummies, as well as state trends. Column (2) adds the logarithm of the real exchange rate interacted the external finance dummy to the specification in the previous column. Column (3) uses the natural logarithm of the transaction value as dependent variable and includes interaction terms between the deregulation indicators and the *continuous* external finance dependence measure. The specification includes source country, mode of entry, and industry dummies, as well as source-country-specific time effects. Column (4) additionally includes the logarithm of the real exchange rate and its interaction with the continuous external finance measure. Robust standard errors clustered at the state level are reported in parentheses. *** denotes significance at the 1 percent level, ** at the 5 percent level, and * at the 10 percent level.

Variable	*Entry Rate*		*ln(Transaction Value)*	
	(1)	(2)	(3)	(4)
Interstate banking	0.03	0.05	-0.28***	-0.27***
	(0.12)	(0.13)	(0.09)	(0.09)
Interstate banking × Ext. fin. dep.	0.34***	0.30**	0.96**	0.93**
	(0.13)	(0.14)	(0.46)	(0.46)
Intrastate branching	-0.06	-0.05	0.10	0.09
	(0.11)	(0.11)	(0.17)	(0.17)
Intrastate branching × Ext. fin. dep.	0.15	0.14	0.50	0.49
	(0.11)	(0.11)	(0.38)	(0.39)
ln(Real exchange rate) × Ext. fin. dep.		-0.003		0.03
		(0.003)		(0.04)
Covariates	Yes	Yes	Yes	Yes
R^2 / Log Pseudo Likelihood	35,594	35,600	0.30	0.30
No. Obs.	1,656	1,656	2,915	2,915
Zero Obs.	593	593	-	-

Panel B: The Impact of Interstate Banking and Branching Deregulation on Average FDI Transaction Value and Entry Rate (No. of New FDI Trans./No. of Multinationals) – Multi-state investors; U.S. Manufacturing Sector, 1977-1994. Similar to Panel A, for the state-level regressions, we categorize the transactions into two groups – those initiated by multi-state foreign investors and those that are initiated by single-state foreign investor. Column (1) presents Tobit results for the entry rate of foreign multinationals (new FDI transactions/number of foreign multinationals) as dependent variable and includes interaction terms between the multi-state investor dummy and the interstate banking and intrastate branching indicators, normalizing the two indicator variables Interstate banking and Intrastate branching by 100. We allow the coefficients on the covariates to be different for the two groups of transactions – those initiated by multi-state foreign investor and those that are not by interacting them with the multi-state investor dummy variable. The specification includes state and year dummies, as well as state trends. Column (2) uses the natural logarithm of the transaction value as dependent variable and includes interaction terms between the deregulation indicators and the multi-state investor dummy. The specification includes source country, mode of entry, and industry dummies, as well as source-country-specific time effects. Robust standard errors clustered at the state level are reported in parentheses. *** denotes significance at the 1 percent level, ** at the 5 percent level, and * at the 10 percent level.

Variable	*Entry Rate*	*ln(Transaction Value)*
	(1)	(2)
Interstate banking	0.07	-0.30*
	(0.10)	(0.16)
Interstate banking × Multi-state Investor	0.18*	-0.05
	(0.10)	(0.24)
Intrastate branching	0.09	0.09
	(0.12)	(0.20)
Intrastate branching × Multi-state Investor	-0.12	-0.09
	(0.10)	(0.22)
Covariates	Yes	Yes
R^2 / Log Pseudo Likelihood	36,350	0.29
No. Obs.	1,656	2,915
Zero Obs.	593	-

Table 9. The Impact of Interstate Banking and Branching Deregulation on Investor's Total Annual Investment in the U.S. Manufacturing Sector, 1977-1994. We specify the following differences-in-differences econometric model with multiple time periods:

$$(1)\quad \ln V\_Total_{ict} = \beta_1 \text{Share\_Interstate\_Bank}_t + \beta_2 \text{Share\_Intrastate\_Branch}_t + \mathbf{Z}_i\alpha + GDP\_Growth_t + Time\_Trend_t + \lambda_c + \varepsilon_{imcstj},$$

where $\ln V\_Total_{ict}$ is the natural logarithm of the real value of the total investment in the U.S. of investor *i* in year *t*. The two variables Share_*Intrastate_Branching*$_t$ and Share_*Interstate_Banking*$_t$ equal to the fraction of U.S. states that have deregulated intrastate bank branching or interstate banking. Vector $\boldsymbol{Z}_i$ includes investor-specific covariates. Our model also controls for GDP growth and includes a time trend and source country fixed effects. Column (1) presents results from a specification that includes GDP growth rate and time trend in addition to the fraction of U.S. states that have deregulated intrastate bank branching or interstate banking. Column (2) presents the results of an augmented specification that includes investor-specific covariates and source country fixed effects. Columns (3)-(4) present the estimates from the specifications in columns (1) and (2) obtained using observations for multiple-transaction investors only. Finally, columns (5) and (6) present results from specifications that additionally include investor-specific fixed effects. Robust standard errors clustered by investor are reported in parentheses. *** denotes significance at the 1 percent level, ** at the 5 percent level, and * at the 10 percent level.

Dependent Variable – *ln*(Total investment value)

Variable	(1)	(2)	(3)	(4)	(5)	(6)
Fraction states with interstate banking	0.84**	1.44***	1.39**	1.92***	2.30***	1.82***
	(0.43)	(0.39)	(0.66)	(0.58)	(0.69)	(0.69)
Fraction states with intrastate branching	0.90	0.58	0.82	-0.07	-0.77	0.04
	(1.02)	(0.97)	(1.58)	(1.51)	(1.66)	(1.62)
GDP growth rate	0.04*	0.04*	0.04	0.05*	0.06*	0.05
	(0.02)	(0.02)	(0.03)	(0.03)	(0.03)	(0.04)
Time trend	-0.04	-0.07**	-0.09	-0.09	-0.04	-0.03
	(0.04)	(0.04)	(0.06)	(0.05)	(0.06)	(0.06)
Multiple-transaction investor		0.89***				
		(0.10)				
Previously invested		0.26		0.29*		-1.16***
		(0.16)		(0.17)		(0.21)
Number of previous investments		0.08**		0.08**		0.22**
		(0.03)		(0.03)		(0.09)
ln(Avg. value of previous investments)		0.21***		0.19***		-0.45***
		(0.04)		(0.05)		(0.06)
Source country fixed effects	No	Yes	No	Yes	No	Yes
Investor fixed effects	No	No	No	No	Yes	Yes
Multiple-transaction investors only	No	No	Yes	Yes	Yes	Yes
R^2	0.03	0.18	0.03	0.17	0.54	0.59
No. Obs.	2,529	2,529	1,086	1,086	1,086	1,086

APPENDIX

Table 3, column (6)

This section describes the technical details of the zero-inflated negative model we adopt to estimate the effect of the banking deregulations on the number of new inbound FDI transactions. Formally, if N_{st} is the number of new FDI transactions in state *s* and year *t*, the zero-inflated negative binomial distribution is given by

$$(1) \qquad P(N_{st}=n)=\begin{cases} p+(1-p)\left(1+\dfrac{\lambda}{\tau}\right)^{-\tau}, & n=0 \\ (1-p)\dfrac{\Gamma(n+\tau)}{n!\Gamma(\tau)}\left(1+\dfrac{\lambda}{\tau}\right)^{-\tau}\left(1+\dfrac{\tau}{\lambda}\right)^{-n}, & n=1,2,... \end{cases}$$

where p, $0 \le p \le 1$, is the mass that the zero-inflated negative binomial distribution assigns to the "extra" zeroes and (1-*p*) is the mass assigned to a negative binomial distribution. The parameter λ is the mean of the negative binomial distribution and τ is a shape parameter that quantifies the amount of overdispersion. The mean and the variance are $E(N_{st})=(1-p)\lambda$ and $Var(N_{st})=(1-p)\lambda(1+p\lambda+\lambda/\tau)$, respectively. Note that the zero-inflated binomial distribution approaches the zero-inflated Poisson as $\tau\to\infty$, and approaches the negative binomial as $p\to 0$. For more details on the zero-inflated negative binomial distribution see Cameron and Trivedi (1998). The zero-inflated regression model relates *p* and λ to the state-level covariates as follows

$$\lambda=\exp(\delta_1 \text{Interstate\_Bank}_{st}+\delta_2 \text{Intrastate\_Branch}_{st}+\mathbf{X_{st}}\rho+\omega_s+\tau_t+\omega_s * Trend_t),$$

and

$$p=\frac{\exp(\pi_1 \text{Interstate\_Bank}_{st}+\pi_2 \text{Intrastate\_Branch}_{st}+\mathbf{X_{st}}\psi+\omega_s+\tau_t)}{1+\exp(\pi_1 \text{Interstate\_Bank}_{st}+\pi_2 \text{Intrastate\_Branch}_{st}+\mathbf{X_{st}}\psi+\omega_s+\tau_t)}.$$ [40]

[40]This specification uses a logistic model to estimate the binary outcome (*n* = 0 or otherwise). The state-specific time trends are included in the determination of positive counts, but excluded from the equation that determines whether the count is zero to accommodate for differences in state-specific growth in the number of new FDI transactions and to lessen the computational burden.

The vector of state-specific, time-varying controls, $\mathbf{X_{st}}$, is the same as the one used in the model for the entry rates and transaction values. As before, we include a full set of state and year dummies. We estimate this zero-inflated negative binomial model using maximum likelihood. The standard errors are adjusted for heteroskedasticity and clustered by state. As we do in the case of the intensive margin and the entry rate specification, we weight by the natural logarithm of the state average manufacturing employment in foreign multinationals over the period 1977-1985. Quantitatively and qualitatively similar results are obtained in unweighted regressions.

APPENDIX TABLES

Table A1. Banking Deregulation Dates.

State	Statewide Branching through M&A Permitted	Interstate Banking Permitted	State	Statewide Branching through M&A Permitted	Interstate Banking Permitted
Alabama	1981	1987	Nebraska	1985	1990
Arizona	Before 1970	1986	Nevada	Before 1970	1985
Arkansas	1994	1989	New Hampshire	1987	1987
California	Before 1970	1987	New Jersey	1977	1986
Colorado	1991	1988	New Mexico	1991	1989
Connecticut	1980	1983	New York	1976	1982
Delaware	Before 1970	1988	North Carolina	Before 1970	1985
Florida	1988	1985	North Dakota	1987	1991
Georgia	1983	1985	Ohio	1979	1985
Idaho	Before 1970	1985	Oklahoma	1988	1987
Illinois	1988	1986	Oregon	1985	1986
Indiana	1989	1986	Pennsylvania	1982	1986
Iowa	1997	1991	Rhode Island	Before 1970	1984
Kansas	1987	1992	South Carolina	Before 1970	1986
Kentucky	1990	1984	South Dakota	Before 1970	1988
Louisiana	1988	1987	Tennessee	1985	1985
Maine	1975	1978	Texas	1988	1987
Maryland	Before 1970	1985	Utah	1981	1984
Massachusetts	1984	1983	Vermont	1970	1988
Michigan	1987	1986	Virginia	1978	1985
Minnesota	1993	1986	Washington	1985	1987
Mississippi	1986	1988	West Virginia	1987	1988
Missouri	1990	1986	Wisconsin	1990	1987
Montana	1990	1993	Wyoming	1988	1987

Source: Amel (1993), Kroszner and Strahan (1999), and Demyanyk et al.(2007).

Table A2. The Impact of Interstate Banking and Branching Deregulation on the Average Foreign Direct Investment Transaction in the U.S. Manufacturing Sector, 1977-1994. We use inverse probability weights to correct for excluding observations with missing transaction value data. All specifications are weighted by the inverse of the predicted probability of selection into the sample, i.e. the probability that data on the value of the transaction are not missing. All specifications include year and state fixed effects. Robust standard errors clustered at the state level are reported in parentheses. *** denotes significance at the 1 percent level, ** at the 5 percent level, and * at the 10 percent level.

Dependent Variable – *ln*(Transaction value)								
Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Interstate banking	-0.36***	-0.34***	-0.35**	-0.28*	-0.27**	-0.26*	-0.23**	-0.27**
	(0.12)	(0.11)	(0.13)	(0.14)	(0.13)	(0.13)	(0.11)	(0.10)
Intrastate branching	0.02	0.06	0.02	0.04	0.02	0.01	0.04	0.07
	(0.13)	(0.13)	(0.17)	(0.16)	(0.16)	(0.17)	(0.18)	(0.17)
ln(GSP)		1.04	1.06	1.04	1.51	1.19	1.91	2.08
		(1.01)	(1.75)	(1.65)	(1.59)	(1.46)	(1.59)	(1.73)
GSP Growth rate		0.83	0.38	0.44	0.27	-0.04	-0.33	-0.88
		(1.79)	(1.85)	(1.89)	(1.89)	(1.78)	(1.91)	(1.95)
GSP Growth rate lag		1.48	1.80	1.38	1.87	2.62	2.78	1.81
		(2.08)	(2.36)	(2.38)	(2.41)	(2.28)	(2.60)	(2.63)
Unemp rate		0.01	-0.03	-0.01	-0.02	-0.02	0.01	0.01
		(0.05)	(0.08)	(0.08)	(0.08)	(0.08)	(0.08)	(0.08)
ln(Wage)		-1.60***	-1.94	-1.58	-2.37	-2.05	-1.58	-1.27
		(0.55)	(1.96)	(1.96)	(1.90)	(1.91)	(1.77)	(1.87)
Corporate tax		-0.12**	-0.16**	-0.14**	-0.16**	-0.16**	-0.20***	-0.19***
		(0.05)	(0.06)	(0.06)	(0.06)	(0.06)	(0.06)	(0.06)
Foreign trade zones		0.02	-0.01	-0.00	-0.01	0.01	0.00	-0.01
		(0.01)	(0.06)	(0.07)	(0.06)	(0.06)	(0.06)	(0.06)
ln(Market potential)		-2.39	-5.22	-4.08	-6.09	-4.70	-6.42	-5.25
		(4.58)	(5.52)	(5.58)	(6.54)	(6.37)	(6.44)	(6.18)
Population density		-0.00	-0.01	-0.01	-0.02	-0.02	-0.02	-0.03
		(0.00)	(0.03)	(0.03)	(0.03)	(0.03)	(0.02)	(0.02)
Multiple-transaction investor								0.54***
								(0.09)
Previously invested								0.22
								(0.14)
Number of previous investments								0.03***
								(0.01)
ln(Avg. value of previous investments)								0.07**
								(0.03)
State Trends	No	No	Yes	Yes	Yes	Yes	Yes	Yes
Source Country Fixed Effects	No	No	No	Yes	Yes	Yes	Yes	Yes
Mode of Entry Fixed Effects	No	No	No	No	Yes	Yes	Yes	Yes
Two-digit Industry Fixed Effects	No	No	No	No	No	Yes	Yes	Yes
Source Country x Year Fixed Effects	No	No	No	No	No	No	Yes	Yes
R^2	0.09	0.09	0.11	0.16	0.19	0.23	0.27	0.31
No. Obs.	2,915	2,915	2,915	2,915	2,915	2,915	2,915	2,915