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 and the resulting improved access to 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. Finally, we demonstrate that following 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 adopting intrastate branching followed by interstate banking deregulation that lifted restrictions on bank expansions. These two types of deregulation have led to higher competition, greater efficiency, and reduction in monopoly power in the banking sector, thereby facilitating access to finance (Jayaratne & Strahan 1996; Jayaratne & Strahan 1998; Cetorelli & Strahan 2006).[1] A number of previous studies have examined the effects of the two banking deregulations and the accompanying reduction in credit constraints 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 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] We show that facilitating access to local bank finance increased the entry rate of foreign multinationals as well as the number of FDI transactions while reducing the average FDI transaction value, i.e. when the cost of borrowing is reduced, smaller FDI transactions are more likely to be completed. Additionally, we provide evidence that as the share of states that improve access to local credit rose, foreign multinationals' total investment in the U.S. grew.

We address the question whether U.S. interstate banking and intrastate branching deregulations have had an impact on FDI flowing into the U.S. Our main hypothesis is that the banking deregulations had a positive impact on FDI activity. We know from the extant literature that multinational firms utilize significant amounts of host country debt financing in their affiliates' capital structure. Such financing is used both for cross-border transactions as well as for the ongoing operations of foreign affiliates. Faccio and Masulis (2005) show that cross-border merger and acquisition deals are more likely to be financed

[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 have also become 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 this study.

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 (Gresik 2001; 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.

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.[3] Therefore, variation in access to external local debt financing could play a significant role for the incidence and the intensity of cross-border transactions. The alternative hypothesis is that improvements in local credit conditions should have little effect on FDI activity since multinational firms have access to their well-developed internal capital markets and have 'deep pockets'.

We employ transaction-level data collected by the International Trade Administration (ITA) of the Department of Commerce to study the impact of state-level banking deregulations on new inward FDI in the U.S. manufacturing sector. 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.[4] 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 facilitating access to 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

[3] 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.
[4] The correlation between the ITA and the Bureau of Economic Analysis (BEA) measures of inbound FDI is 0.86 between 1979 and 1990 (Klein & Rosengren 1994).

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 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.

We show that whereas the deregulation of intrastate branching did not have any significant effect on FDI inflows along the extensive or the intensive margin[5], 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. We also find that when 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 17 percent increase in the number of inbound FDI transactions, translating to 1.18 new FDI transactions per year, and an increase in the entry rate of foreign multinationals from 1.15 to 1.63 percent.

Investigating the impact of banking deregulation along the intensive margin, we find that the average value of foreign transactions decreased by approximately 22.1 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

[5] The extensive margin refers to the incidence of FDI or the number of transactions and the entry rate while the intensive margin refers to the intensity of FDI activity or transaction values.

indicate that with cheaper external finance, foreign firms were able to undertake projects of smaller value, which became profitable with lower borrowing costs. 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 to 18 percent.

To illuminate the mechanisms behind the effect of banking deregulation on the incidence and the intensity of FDI, we extend our study in two directions. First, we make the distinction between multiple and single transaction investors. We hypothesize that multiple transaction investors are more likely to avail themselves of local bank finance since they have prior exposure to the U.S. market, which first-time investors lack. 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 local finance were important for inbound FDI activity, we would anticipate the effects of banking deregulation to be more pronounced in industries that are more reliant on external finance.

We 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. To establish how transaction values vary depending on investor experience, we split the sample into single versus multiple transaction investors. We find that transaction values for foreign firms engaging in multiple investments declined by 35.6 percent following interstate banking deregulation, while single-transaction investors experienced declines of similar magnitudes to the overall sample. Our results also suggest that for a given investor, transaction size declined with the number of previous transactions, indicating that, on average, investors undertake major projects upon entering the U.S. market and subsequently make smaller adjustments. Our results are consistent with interstate banking deregulation creating economies of scale for multiple transaction investors as they were enabled to consolidate their banking relationships across 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 multinationals was far more pronounced in the *more* external finance dependent industries. Hence, by facilitating access to credit, interstate banking deregulation allowed a larger number of firms that rely on external finance to invest in the U.S. Along the intensive margin we find that while on average transaction values decreased following the interstate banking deregulation, transaction values in sectors *more* dependent on external finance increased vis-à-vis transaction values in *less* 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. Possibly, national banks have a comparative advantage in evaluating foreign investment projects and multi-state banks have better technology to serve multinational firms investing in the U.S. relative to single-state banks.

Our study contributes to a small, but growing literature assessing the effects of credit constraints on international economic activity (Manova 2008; Amiti & Weinstein 2011; Chor & Manova 2012). 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.

The link between access to bank finance and real economic activity has been explored at length in the domestic context (Levine 2005). Cetorelli & Strahan (2006) and Kerr & Nanda (2009) have shown that firm entry and entrepreneurship among domestic firms react positively to banking deregulation. Additionally, Michalski and Ors (2012) have shown that bilateral trade increased in state-pairs that liberalized their banking systems. 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.

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. The 1956 Douglas Amendment of the Bank Holding Company Act prohibited bank holding companies from acquiring banks in other states unless state regulations permitted such transactions. Aside from the handful of grandfathered multistate holding companies, this effectively banned interstate bank mergers and acquisitions since no state allowed such cross-state transactions. 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 Table 1). The Garn-St. Germain Act of 1982 further amended the Bank Holding Company Act to allow any bank holding company, regardless of its state, to acquire failed banks (Jayaratne & Strahan 1996). However, it was not until the Riegle-Neal Interstate Banking and Branching Efficiency Act of 1994 that interstate banking was deregulated nationwide, unless 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).[6]

Similarly, until the 1970s only a handful of states allowed unrestricted within state branching. The majority of states either explicitly prohibited or severely limited bank branching activity (Jayaratne & Strahan 1996), although banks could effectively branch by adopting a multi-bank holding company organizational form (Kerr & Nanda 2009). 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. Branching through M&As allowed multi-bank holding companies to transform subsidiaries into branches, as well as to acquire branches. Most states permitted de novo branching (the set up of brand new branches) at a later stage. 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.

[6] 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).

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 of the preponderance of credit card banks in these states (Black & Strahan 2002; Berger *et al.* 2012).

We present data on U.S. commercial bank lending to foreign firms juxtaposed with foreign investment activity at the state level in Table 2. The first two columns of Table 2 present Federal Depository Insurance Corporation (FDIC) data from 1993 commercial and industrial loans issued to foreign borrowers (non-U.S. addressees) by U.S. depository institutions along with their percentage in overall commercial and industrial lending.[7] These data are drawn from the FDIC Statistics on Depository Institutions Report and are aggregated at the state level. The numbers represent commercial and industrial lending to foreign borrowers made by all FDIC-insured institutions, comprising both commercial banks and savings institutions, which encompass the population of institutions that are affected by state banking deregulation laws. Commercial and industrial loans consist of loans issued to firms predominantly in the manufacturing industry and exclude real estate and farm loans. Column (3) reports the value of all manufacturing transactions with non-missing values in 1993 in our sample while column (4) reports the total number of manufacturing transactions, both with missing and non-missing transaction values. Column (5) presents the 1993 gross state product expressed in 1983 dollars. There is a high correlation (0.7) between the value of commercial and industrial loans to non-U.S. addressees reported in column (2) and the total value of manufacturing transactions reported in column (3), indicative of a tight link between local bank financing and state-level FDI activity.

[7] The choice of year is dictated by FDIC data availability and the 1994 passage of the Riegle-Neal Interstate Banking and Branching Efficiency Act.

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.[8] 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 origin), location of the investment (state) and the year the transaction was completed.[9] 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 (see Table 3, Panel B).

To assess the effect of increased access to 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,

[8] After 1994, only the Bureau of Economic Analysis (BEA) collects such data, however their data are confidential and not publicly available.

[9] 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.

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 in 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 on FDI (Keller & Levinson 2002).

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. Specifically, Klein and Rosengren (2002) report that the correlation between the BEA measure of inward FDI and the ITA measure of total inward FDI between 1979 and 1990 is 0.86. 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 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 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.[10] 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 Table A1.

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. $\text{Market Potential}_{\text{st}} = \sum_{\text{n}\neq\text{s}} \frac{\text{GSP}_{\text{nt}}}{\text{Distance}_{\text{ns}}}$); and population density calculated as state population divided by total land area.

[10] The entry rate is defined as the number of new inbound FDI transactions as a fraction of the total number of existing foreign multinationals. Because data on the number of multinationals in the manufacturing sector 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.

Summary statistics for all variables included in our analysis are presented in Panels A and B of Table 3. 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 a given state, 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 $70.7 million (1983 dollars), but there is considerable variation – the smallest transaction is only $67,500 while the largest is over $7 billion.

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 entry rate as the number of new inbound FDI transactions normalized by the total number of multinationals present in each state. 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) $\text{Entry\_Rate}^*_{\text{st}} = \alpha_1 \text{Interstate\_Bank}_{st} + \alpha_2 \text{Intrastate\_Branch}_{st} + \mathbf{X_{st}}\mu + \omega_s + \tau_t + \omega_s * Trend_t + \kappa_{st}$

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

where $\text{Entry\_Rate}^*_{\text{st}}$ is the underlying latent variable, which is not observed, and it satisfies the classical linear model assumptions and $\text{Entry\_Rate}_{\text{st}}$ 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 10 above). Equations (1) and (2) above imply that the observed variable, $\text{Entry\_Rate}_{\text{st}}$, equals $\text{Entry\_Rate}^*_{\text{st}}$ when $\text{Entry\_Rate}^*_{\text{st}} \geq 0$, and $\text{Entry\_Rate}_{\text{st}} = 0$ when $\text{Entry\_Rate}^*_{\text{st}} < 0$.

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).[11]

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 number and rate of inward 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 direct 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).

We estimate the Tobit model using maximum likelihood estimation. The standard errors are adjusted for heteroskedasticity and are clustered by state. We weight all of the empirical specifications by the log of the average state manufacturing employment in foreign multinationals over the period 1977-1985 (see, for example, Kerr and Nanada 2009).[12] 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, or when we weight by the average state manufacturing employment in foreign multinationals.

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[11] Haung (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).

[12] 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 10).

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 and Wooldridge 2001 for details), 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-years 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. We present the details of the zero-inflated negative binomial model we estimate in the Technical Appendix.

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:

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

where $\log V_{imcstj}$ 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.

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 potentially 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. 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 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 log of the average state manufacturing employment in foreign multinationals over the period 1977-1985. Qualitatively and quantitatively similar results are obtained in unweighted regressions.

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 deregulation affect the total value of foreign investments undertaken by multinationals that invest multiple times in the U.S.

Before we discuss the formal results from our econometric models that are rooted in pre-post deregulation comparisons, we offer a visual illustration of the impacts. To this end, we estimate dynamic models for the entry rate and transaction values, that are similar to those laid out in the previous section (equations (1), (2) and (3)), but include 20 separate indicator variables, instead of only one pre-post indicator, tracing over time the passage of each of the two banking deregulations:

$$\text{(5)}\quad Outcome_{st} = \sum_{i=-10}^{10} \beta_{1,t+i} \text{Interstate\_Bank}_{s,t+i} + \sum_{i=-10}^{10} \beta_{2,t+i} \text{Intrastate\_Branch}_{s,t+i} + \mathbf{X_{st}}\gamma + \omega_s + \tau_t + \omega_s * Trend_t + \nu_{st}$$

The dummy variables take on a value of one in the i^{th} year before or after the deregulation and are zero otherwise.[13] The -10 and +10 year endpoints include all years earlier and later than the 20-year window. An indicator for the year prior to the deregulation is not included, so that the estimated coefficients measure the year-by-year dynamics of the entry rates or the (log of) transaction values relative to the year prior to the reform year. The estimated series for $\beta_{1,t+i}$ and $\beta_{2,t+i}$ are plotted in Figures 1 and 2. The graphs show that lead effects for both deregulations are small and statistically insignificant. After deregulation, no changes in the average transaction value or the entry rate are evident for the intrastate branching deregulation; however, there is a pattern of an increasing entry rate and a declining average transaction value following the interstate banking deregulation.

[13] The model for the average transaction value additionally includes source country, mode of entry, and industry dummies.

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

In Table 4, we present 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 the last four columns we consider only the number of new transactions with recorded transaction values. In both cases, the denominator is the total number of existing multinationals in operation in that state and year. Columns (1) - (3) and (5) - (7) report the results from the Tobit model, whereas columns (4) and (8) report 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 (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 3, 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 finding suggests that the increased availability of credit as a result of the interstate banking deregulation allows a higher incidence of FDI transactions. Our finding of the importance of *host* country financing 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.[14] 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 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.

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 4 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 all specifications in columns (5) - (8) are 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

[14] 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 4, 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.

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 4 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 3) 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).

Next, we consider the impact of the two banking deregulations on the number of new inbound FDI transactions. Table 5 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 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 5. Column (1) presents the results from the baseline specification without covariates for the total number of new inbound FDI transactions. Consistent with our findings from the specifications for the entry rate (see Table 4), 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 0.30 (with a standard error of 0.10) on interstate banking implies that the number of new inbound FDI transactions increased by 35 percent following the adoption of the interstate banking deregulation.[15] Since the average number of new transactions (including those with missing transaction values) is 6.93 (see Table 3, Panel A), a 35 percent increase implies that the average state recorded 2.43 additional transactions annually after interstate banking deregulation occurred.

When we augment the baseline specification with additional covariates and state trends, the coefficient on interstate banking declines to 0.17 (with a standard error of 0.09), but remains both statistically and economically significant. The estimate on interstate banking in our preferred specification in column (3) suggests that in states, which deregulated interstate banking, the number of new foreign investment transactions rose by 17 percent, or equivalently, such states experienced an increase of about 1.18 transactions annually. Overall, the results from the count models are consistent with the evidence from the entry rate regressions, suggesting that both the entry rate of multinationals and the number of new inbound FDI transactions increase following the adoption of interstate banking deregulation.

The last two columns of Table 5 present the estimates of the impact of interstate banking and intrastate branching deregulations on the number of transactions for single-transaction and for multiple-transaction investors, respectively. The estimates for both groups are positive and very similar to the overall full-sample estimates in column (3). The coefficient on the interstate banking deregulation indicator is statistically significant for the multiple-transaction investor group, showing deregulation led such investors to complete greater number of transactions. This result suggests that multiple-transaction investors, which are more likely to have prior exposure to U.S. banks compared with single-transaction investors, are more prone to be affected by changes in access to local bank finance.

Lastly, we analyze whether the impact of banking deregulation varies with the degree of external finance dependence. We would expect the impact of the two banking deregulations to be more pronounced in sectors that are more reliant on external finance (Rajan & Zingales 1998). To that end, we

[15] Because the indicator variable only changes discontinuously, the effect of the interstate banking deregulation is calculated as ($e^{0.30}$-1) = 0.35. For estimated coefficients that are small in magnitude, this procedure makes little difference.

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).[16] 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 then construct an external finance dependence dummy 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 the interstate banking and intrastate branching indicators in our preferred Tobit specification with covariates and state trends for the entry rates.[17]

In column (1) 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 alleviating credit constraints, banking deregulation allowed a larger number of

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[16] 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.

[17] 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.

firms 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 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.

5.2 Impact on the Average FDI Transaction Value

Table 6 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 6 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 indicator, suggesting that the average FDI transaction value declined by 29.5 percent following the adoption of the interstate banking deregulation.[18] This finding suggests that the increased availability of credit as a result of the interstate banking deregulation created economies of scale allowing smaller FDI transactions to take place because of the lower cost of banking. This, in turn, lowered the average value of foreign direct investment transactions in the U.S. We show more evidence consistent with this economies of scale argument by examining the investment patterns of multiple transaction investors doing business across multiple states. 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.

In the second column of Table 6, 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

[18] 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 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 6, 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 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 3 for the list of countries).[19] These investor country-by-year effects control for 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 the exhaustive specification in column (7), we find the coefficient on interstate banking deregulation to be -0.20 (with a standard error of 0.09). This estimate suggests that by increasing the availability of credit, interstate banking deregulation allowed smaller transactions to be financed and led to an 18.1 percent decline in the average foreign transaction value.

In the last column of Table 6, we 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 source-country year effects. The coefficient on interstate banking increases in magnitude to -0.25 (with a standard deviation of 0.09), and is significant at the 1 percent level, suggesting that the passage of the interstate banking deregulation lead to a 22.1 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 78.6 (= $(e^{0.58}-1)*100$) percent larger. The other variables pertaining to previous investment behavior are all positive, with the number of previous investments and the logarithm of the average value of previous investments being statistically significant.

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 is not missing at 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

[19] 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 time effects.

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 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 A1. The estimated effects of both banking deregulations are very similar to those reported in Table 6, where we do not correct for missing values.

In Table 7, we further analyze how transaction values change with previous investment history by carrying out a subsample analysis. We first separate the single-transaction investors in our sample from the multiple-transaction investors, and estimate the augmented specification (3), which corresponds to column (8) of Table 6, on these two subsamples. Columns (1) and (2) of Table 7 present the results for the single and multiple-transaction investors separately. 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 (compared to a decrease of 22.1 percent for the full sample).

Turning to the investor- and transaction-specific covariates in column (3) of Table 7, we find that the coefficients on the indicator variable for having previously invested and the logarithm of the average value of previous investments are now negative and highly significant, unlike their counterparts in column (8) of Table 6. This suggests that for a given investor, transaction size declines with the number of previous transactions, indicating that, on average, investors undertake major projects upon entering the U.S. market and subsequently make smaller adjustments. The evidence is also consistent with the notion that interstate banking deregulation created economies of scale for multiple transaction investors by lowering the cost of banking across state borders and allowing investors to make smaller investments on average.

In columns (4) and (5) of Table 7, we check the robustness of the results for the 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) to column (2) of Table 7 and column (5) to column (3) of Table 7.

Finally, we consider how the interstate banking and branching deregulations impact the average foreign transaction value taking place in sectors more reliant on external finance versus those in sectors that are less reliant on external finance. To formally test if the effect of banking deregulation on transaction values changes with the need for external finance, we include interaction terms between the deregulation indicators and the continuous external finance measure of Cetorelli and Strahan (1998) in equation (3). Column (2) of Table 8 presents these results. The main effect of interstate banking remains very similar to the findings in Table 6, 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 local finance improved. As in all previous specifications, neither the interaction term nor the main effect of the intrastate branching deregulation is significant.

5.3 Impact on Total Investment Volume of Foreign Multinationals

Next, 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, especially those that invest more than once, 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, whether the banking deregulations led foreign firms to increase the total value of their investments in the U.S. is ambiguous.

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 summing up all transactions (across different states, industries, and modes of entry) in a year completed by a given investor. 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.:

$$(6)\ \log V\_Total_{ict} = \beta_1 \text{Share\_Interstate\_Bank}_t + \beta_2 \text{Share\_Intrastate\_Branch}_t + \mathbf{Z}_{it}\alpha + $$

$$+ GDP\_Growth_t + Time\_Trend_t + \lambda_c + \varepsilon_{ict}.$$

The dependent variable, $\log V\_Total_{ict}$, 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. 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 impacts of access to credit in two important ways. First, we provide direct micro-level evidence from the U.S. that improved access to *local* finance 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. We employ information on both the value of each transaction (the intensive margin) and the number of transactions (the extensive margin) to evaluate how improved access to credit affects inbound FDI and the entry rate of foreign multinationals. Second, we extend the current literature by providing estimates of the impact of changes in access to *local* finance on foreign firms investing in the U.S. To this end, we employ cross-state variation in the timing of two financial deregulations, intrastate branching and interstate banking, that increased local banking competition and improved access to local credit. Specifically, 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), that 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 17 percent increase in new FDI projects, which translates into an additional 1.18 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 local credit improves as a result of increased bank competition, foreign firms are able to undertake smaller inbound FDI projects.

Taken together, these results imply that by facilitating access to local external 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 foreign 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 mechanism behind the effect of the interstate banking deregulation, we check how the impact differs between single-transaction investors and multiple-transaction investors, who are more likely to avail themselves of local bank finance since they have prior exposure to the U.S. market, which first-time investors lack. Consistent with this hypothesis, we find that while both single and multiple transaction investors experience an increase in the number of completed transactions and entry rates, the effect is larger in magnitude for the multiple transaction investors. We also provide direct evidence of the importance of the local finance channel for FDI by comparing the impact of banking deregulation on foreign transactions in sectors that rely more heavily on external finance relative to those in sectors, which are less reliant on external finance (Rajan and Zingales 1998, Cetorelli and Strahan 1998). If access to local finance were important for inbound FDI activity, the effects of the interstate banking deregulation should be more pronounced in industries that are more reliant on external finance, which is what we find. Both transaction values and entry rates for FDI in industries that are *more* dependent on external finance increase after the adoption of interstate banking deregulation relative to industries that are less reliant on external finance.

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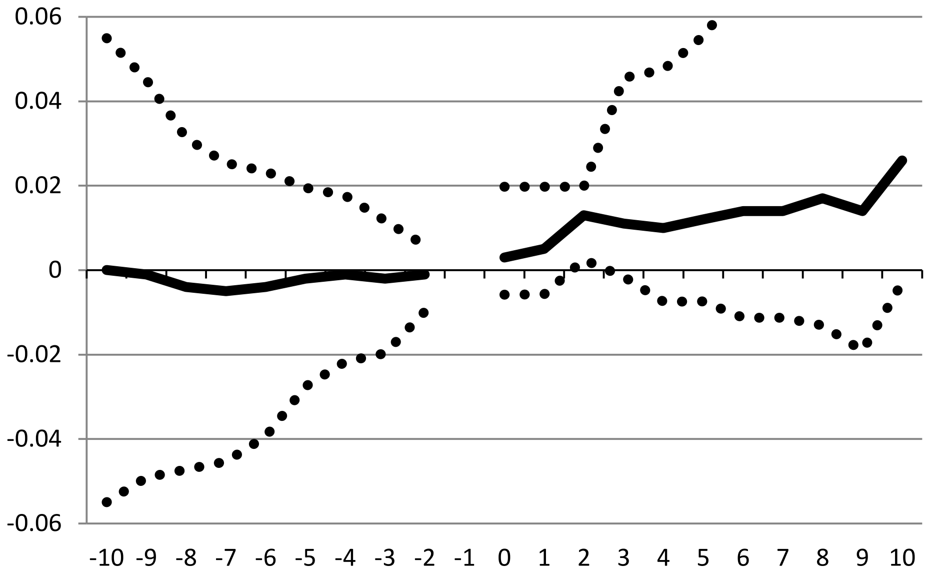
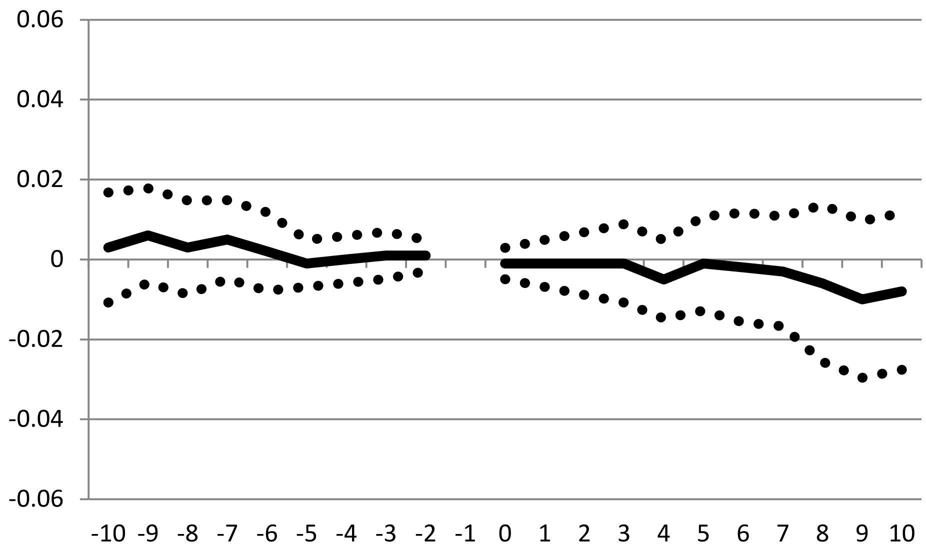
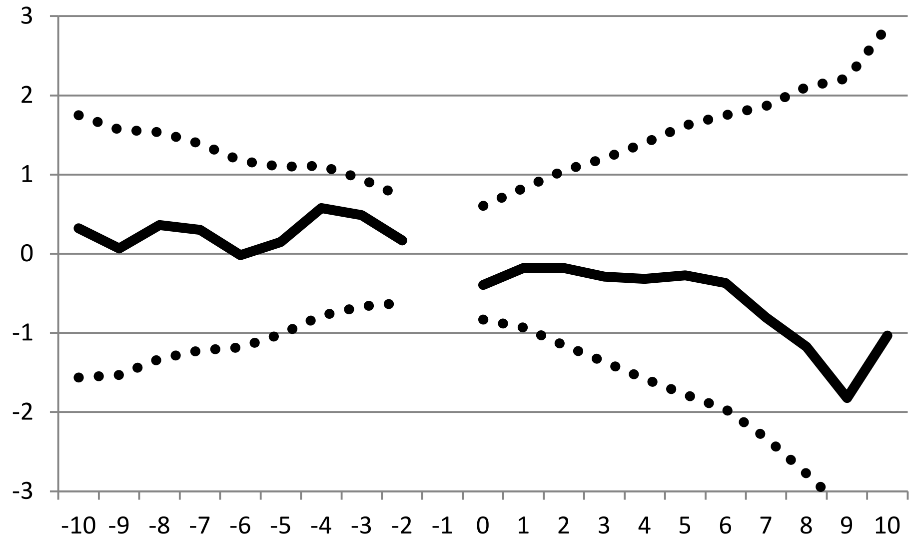


Interstate Banking

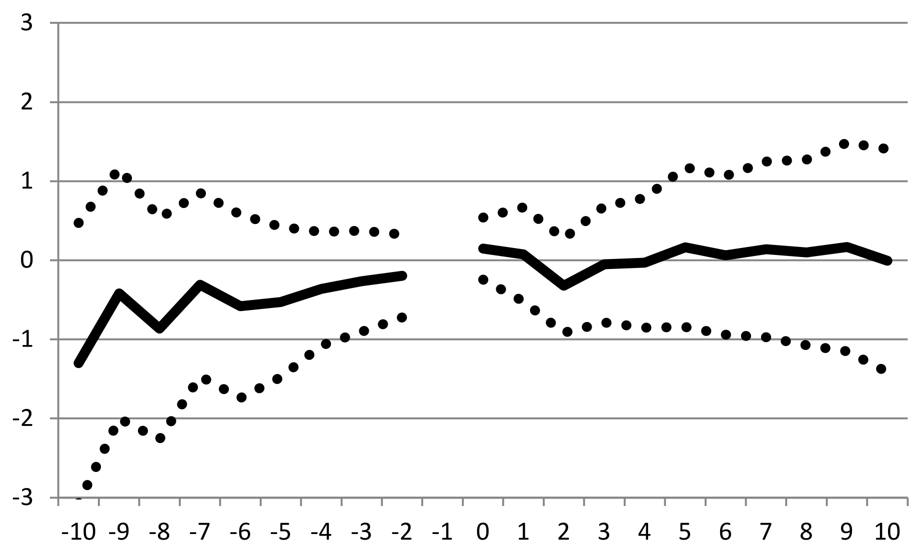


Intrastate Branching

Figure 1. 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. The standard error bars in the top panel are linearly transformed as follows: $\frac{\text{95\% Confidence Interval}}{0.5}$.



Interstate Banking



Intrastate Branching

Figure 2. 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. The standard error bars in the top panel are linearly transformed as follows: $\pm 0.5 + \frac{\text{95\% Confidence Interval}}{3.5}$.

Table 1. 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 2. Year 1993 bank commercial and industrial loans juxtaposed with FDI volume and counts. Columns (1) and (2) are drawn from the Federal Depository Insurance Corporation (FDIC) 1993 Statistics on Depository Institutions Report, Net Loans and Leases. Column (1) reports commercial and industrial loans issued to non-U.S. addressees by U.S. depository institutions. Column (2) reports the percentage of commercial and industrial loans to non-U.S. addressees in overall commercial and industrial lending. Column (3) reports the value of all manufacturing transactions with non-missing values in 1993 in our sample while column (4) reports the total number of manufacturing transactions. Column (5) presents the 1993 gross state product. All dollar variables expressed in millions of 1983 USD.

State	Commercial and Industrial Loans to Non-U.S. Addressees (1983 U.S. $, millions)	Percentage of Non-U.S. Addressees to Overall Commercial and Industrial Loans	Value of Manufacturing FDI Transactions (1983 U.S. $, millions)	Number of Manufacturing FDI Transactions	Gross State Product (1983 U.S. $, millions)
	(1)	(2)	(3)	(4)	(5)
Alabama	1.25	0.03	0.00	0	58,521
Arizona	10.91	0.58	0.00	0	61,799
Arkansas	0.07	0.00	67.82	1	32,918
California	8,050.86	21.23	1,430.87	32	571,936
Colorado	0.36	0.02	0.00	2	66,200
Connecticut	2.19	0.05	207.61	2	74,519
Delaware	115.34	3.29	0.00	0	16,269
Florida	226.29	2.52	10.38	8	210,660
Georgia	60.95	0.59	51.90	7	118,795
Idaho	0.00	0.00	0.00	0	15,986
Illinois	920.40	3.75	224.22	7	221,508
Indiana	0.78	0.01	55.02	4	91,424
Iowa	3.29	0.13	0.00	1	43,866
Kansas	0.01	0.00	245.40	5	40,798
Kentucky	4.84	0.13	218.96	4	56,483
Louisiana	2.03	0.08	156.40	3	66,343
Maine	0.29	0.02	0.00	0	17,441
Maryland	142.21	3.66	100.35	2	86,862
Massachusetts	3,346.32	21.81	1,208.30	5	120,088

Michigan	862.04	6.15	54.46	6	156,656
Minnesota	8.47	0.13	70.93	1	81,624
Mississippi	0.00	0.00	55.36	2	32,761
Missouri	6.62	0.11	116.96	4	82,288
Montana	0.00	0.00	0.00	0	11,110
Nebraska	0.01	0.00	0.00	0	27,234
Nevada	0.00	0.00	56.75	3	27,709
New Hampshire	1.15	0.19	0.00	0	18,936
New Jersey	11.92	0.14	37.37	7	166,359
New Mexico	0.01	0.00	0.00	0	25,520
New York	38,903.10	42.83	2,481.87	16	371,975
North Carolina	345.98	2.96	291.00	15	116,267
North Dakota	0.00	0.00	0.00	0	9,030
Ohio	11.39	0.09	324.43	12	179,662
Oklahoma	1.16	0.05	8.44	1	45,365
Oregon	3.71	0.11	53.08	3	48,034
Pennsylvania	647.57	2.75	227.68	8	194,951
Rhode Island	0.06	0.00	0.00	0	16,309
South Carolina	0.00	0.00	69.07	8	52,704
South Dakota	0.00	0.00	0.00	0	11,145
Tennessee	1.71	0.03	66.78	6	82,746
Texas	420.13	2.30	2,611.49	16	307,111
Utah	0.00	0.00	0.00	0	26,734
Vermont	0.49	0.09	0.00	0	9,015
Virginia	8.92	0.15	23.18	7	116,047
Washington	1,137.34	20.44	1.04	2	98,616
West Virginia	0.00	0.00	0.00	1	22,063
Wisconsin	12.74	0.21	146.92	2	83,062
Wyoming	0.00	0.00	0.00	0	9,544

Table 3. 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
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 3. 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

Table 4. The Impact of Interstate Banking and Branching Deregulation on the Entry Rate of Foreign Multinationals in the U.S. Manufacturing Sector, 1977-1994. The table presents results for the entry rate of foreign multinationals using Tobit and OLS estimators. We estimate the following Tobit Model:

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

$$\text{Entry\_Rate}_{st} = \max\{0, \text{Entry\_Rate}^*_{st}\},$$

The dependent variable 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 columns (5)-(8) the observations with missing values are excluded when calculating the entry rate. 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. Columns (1) and (5) present results from a Tobit specification that includes the deregulation indicators along with a full set of state and year fixed effects. Columns (2) and (6) present the results of an expanded Tobit specification that includes a set of time-varying, state-specific determinants of foreign investment transactions. Columns (3) and (7) augment the specification in columns (2) and (6) with state-specific time trends. Columns (4) and (8) present the results for the augmented specification estimated using OLS. 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)	(7)	(8)
	including observations with missing values				*excluding observations with missing values*			
Interstate banking/100	0.48***	0.48***	0.38**	0.45**	0.43***	0.44***	0.37***	0.38***
	(0.18)	(0.18)	(0.18)	(0.18)	(0.13)	(0.14)	(0.14)	(0.13)
Intrastate branching/100	0.25	0.06	0.07	0.04	0.09	-0.01	-0.04	-0.04
	(0.16)	(0.14)	(0.19)	(0.19)	(0.10)	(0.11)	(0.17)	(0.15)
ln(GSP)		-0.03***	-0.01	0.01		-0.02***	0.00	0.01
		(0.01)	(0.02)	(0.01)		(0.01)	(0.01)	(0.01)
GSP Growth rate		-0.03	-0.06***	-0.04**		-0.02	-0.03**	-0.02**
		(0.02)	(0.02)	(0.02)		(0.01)	(0.01)	(0.01)
GSP Growth rate lag		0.05**	0.03	0.02		0.03	0.02	0.02
		(0.02)	(0.02)	(0.01)		(0.02)	(0.02)	(0.01)
Unemp rate/100		0.03	-0.07	-0.10**		0.02	0.02	-0.02
		(0.05)	(0.06)	(0.05)		(0.04)	(0.04)	(0.03)
ln(Wage)		0.04***	0.01	-0.01		0.02***	0.01	-0.01
		(0.01)	(0.016)	(0.01)		(0.01)	(0.01)	(0.01)
Corporate tax/100		-0.03	-0.18	-0.19**		0.02	0.02	-0.03
		(0.09)	(0.13)	(0.10)		(0.06)	(0.10)	(0.07)
Foreign trade zone/100		-0.00	0.06	0.07		-0.01	0.02	0.02
		(0.02)	(0.08)	(0.08)		(0.01)	(0.07)	(0.06)
ln(Market potential)		-0.04	-0.05	-0.11**		-0.03	0.03	-0.02
		(0.06)	(0.07)	(0.05)		(0.04)	(0.05)	(0.03)
Population density/100		0.01	-0.03	-0.03		0.00	-0.04**	-0.04*
		(0.01)	(0.03)	(0.02)		(0.01)	(0.02)	(0.02)
State trends	No	No	Yes	Yes	No	No	Yes	Yes
Model specification	Tobit	Tobit	Tobit	OLS	Tobit	Tobit	Tobit	OLS
Log Likelihood/ R^2	19,636	19,962	20,442	0.60	19,024	19,223	19,663	0.52
No. Obs.	828	828	828	828	828	828	828	828
Zero Obs.	183	183	183	183	262	262	262	262

Table 5. The Impact of Interstate Banking and Branching Deregulation on the Number of New FDI Transactions in the U.S. Manufacturing Sector, 1977-1994. The table presents results from a zero-inflated negative binomial model for the number of new FDI transactions *N*, in each state *s* (including those with missing values):

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

$$\lambda=\exp(\delta_1\text{Interstate\_Bank}_{st}+\delta_2\text{Intrastate\_Branch}_{st}+\mathbf{X_{st}}\rho+\omega_s+\tau_t+\omega_s*Trend_t)$$
$$p=\frac{\exp(\pi_1\text{Interstate\_Bank}_{st}+\pi_2\text{Intrastate\_Branch}_{st}+\mathbf{X_{st}}\psi+\omega_s+\tau_t)}{1+\exp(\pi_1\text{Interstate\_Bank}_{st}+\pi_2\text{Intrastate\_Branch}_{st}+\mathbf{X_{st}}\psi+\omega_s+\tau_t)}.$$

The specifications in columns (1)-(3) use a dependent variable constructed based on all new FDI transactions. For the specifications in columns (4) and (5), we construct the dependent variables using the number of transactions for single-transaction investors and for multiple-transaction investors, respectively. Column (1) presents the 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. Column (3) further includes state-specific time-trends. Columns (4) and (5) present the results from the augmented specification with covariates and state-specific time trends for the single-transaction investor and the multiple-transaction investor subsamples. 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 – No. of New FDI Transactions

Variable	(1)	(2)	(3)	(4)	(5)
		All transactions		*Single-transaction investors*	*Multiple-transaction investors*
Interstate banking	0.30***	0.27**	0.17**	0.15	0.18**
	(0.10)	(0.11)	(0.09)	(0.14)	(0.08)
Intrastate branching	0.06	0.09	0.03	0.08	-0.01
	(0.09)	(0.10)	(0.11)	(0.14)	(0.12)
ln(GSP)		-0.01	-0.01	-0.01	-0.01
		(0.87)	(1.00)	(1.51)	(1.29)
GSP Growth rate		-0.59	-1.21	-0.47	-1.00
		(1.08)	(0.91)	(1.73)	(1.11)
GSP Growth rate lag		-1.23	-0.61	-2.33	0.01
		(0.90)	(0.77)	(1.43)	(1.19)
Unemp rate		-0.06	-0.06	-0.05	-0.05
		(0.05)	(0.04)	(0.07)	(0.05)
ln(Wage)		-0.02	-0.02	-0.01	-0.02
		(1.10)	(1.03)	(2.16)	(1.11)
Corporate tax		0.00	-0.00	0.00	-0.00
		(0.06)	(0.04)	(0.06)	(0.05)
Foreign trade zone		0.01	0.03	-0.00	0.03
		(0.02)	(0.05)	(0.05)	(0.05)
ln(Market potential)		-0.03	-0.03	-0.02	-0.03
		(3.29)	(2.78)	(3.27)	(4.70)
Population density		0.00	0.00	0.00	0.00
		(0.01)	(0.02)	(0.014)	(0.03)
State trends	No	No	Yes	Yes	Yes
Log Likelihood	-16,419	-16,303	-15,410	-11,294	-12,921
No. Obs.	828	828	828	828	828
Zero Obs.	183	183	183	301	256

Table 6. The Impact of Interstate Banking and Branching Deregulation on Average FDI Transaction Value in the U.S. Manufacturing Sector, 1977-1994. We specify the following differences-in-differences econometric model with multiple time periods:

$$(1)\ \log V_{imcstj} = \beta_1 \text{Interstate\_Bank}_{st} + \beta_2 \text{Intrastate\_Branch}_{st} + \mathbf{X_{st}}\gamma + \mathbf{Z}_i\alpha + \omega_s + \tau_t + \delta_m + \lambda_c + \psi_j + \omega_s * Trend_t + \varepsilon_{imcstj},$$

where *log* 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 vector $\boldsymbol{X}_{imcst}$ includes a host of time-varying, state-specific control variables that are likely to affect incoming FDI and may be correlated with bank branching regulations. Vector $\mathbf{Z}_i$ includes investor and transaction-specific covariates. Our model features state fixed effects ω_s, year fixed effects τ_t, mode of entry fixed effects δ_m, source country fixed effects λ_c, two-digit SIC industry fixed effects φ_j, state-specific time trends, $\omega_s*Trend_t$. ε_{imcst} denotes the residual. We use 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. In column (7), we further include source-country-specific year effects in addition to the full set of fixed effects. Column (8) reports results with investor- and transaction-specific covariates, in addition to the state-specific covariates, a full set of fixed effects and source country 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.

Dependent Variable – *ln*(Transaction value)

Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Interstate banking	-0.35***	-0.32***	-0.33**	-0.29**	-0.27**	-0.25**	-0.20**	-0.25***
	(0.12)	(0.11)	(0.13)	(0.14)	(0.13)	(0.13)	(0.09)	(0.09)
Intrastate branching	0.02	0.05	0.04	0.06	0.05	0.04	0.08	0.09
	(0.12)	(0.12)	(0.16)	(0.16)	(0.15)	(0.16)	(0.17)	(0.17)
ln(GSP)		0.93	0.58	0.68	1.07	0.90	1.57	1.74
		(0.93)	(1.70)	(1.61)	(1.52)	(1.39)	(1.50)	(1.64)
GSP Growth rate		1.25	1.04	0.96	0.72	0.59	0.07	-0.96
		(1.68)	(1.83)	(1.88)	(1.80)	(1.73)	(1.87)	(1.87)
GSP Growth rate lag		2.20	2.75	2.12	2.38	3.26	3.39	2.95
		(2.13)	(2.39)	(2.43)	(2.47)	(2.31)	(2.60)	(2.50)
Unemp rate		0.03	-0.02	-0.00	-0.01	-0.01	0.02	0.01
		(0.05)	(0.07)	(0.07)	(0.07)	(0.07)	(0.07)	(0.07)
ln(Wage)		-1.50***	-2.47	-2.09	-2.74	-2.49	-1.89	-1.63
		(0.51)	(1.81)	(1.81)	(1.78)	(1.78)	(1.64)	(1.75)
Corporate tax		-0.13***	-0.16***	-0.15***	-0.17***	-0.18***	-0.22***	-0.20***
		(0.05)	(0.06)	(0.05)	(0.06)	(0.05)	(0.06)	(0.06)
Foreign trade zones		0.03**	-0.00	0.00	-0.01	0.01	0.00	-0.01
		(0.01)	(0.06)	(0.07)	(0.06)	(0.06)	(0.07)	(0.06)
ln(Market potential)		-0.70	-3.22	-2.28	-3.68	-2.43	-3.56	-4.38
		(4.30)	(5.34)	(5.41)	(6.17)	(5.98)	(6.31)	(6.52)
Population density		-0.00	0.00	0.00	-0.00	-0.01	-0.01	-0.02
		(0.00)	(0.03)	(0.03)	(0.03)	(0.03)	(0.02)	(0.02)
Multiple-transaction investor								0.58***
								(0.08)
Previously invested								0.16
								(0.13)
Number of previous investments								0.03**
								(0.01)
ln(Avg. value of previous investments)								0.11***
								(0.02)
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 × year fixed effects	No	No	No	No	No	No	Yes	Yes
R^2	0.079	0.085	0.099	0.148	0.179	0.218	0.255	0.295
No. Obs.	2,915	2,915	2,915	2,915	2,915	2,915	2,915	2,915

Table 7. 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 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 Results; U.S. Manufacturing Sector, 1977-1994. 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) 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. 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	(1)	(2)
	Entry Rate	*ln(Transaction Value)*
Interstate banking	0.03	-0.28***
	(0.12)	(0.09)
Interstate banking × Ext. fin. dep.	0.344***	0.96**
	(0.13)	(0.46)
Intrastate branching	-0.06	0.10
	(0.109)	(0.17)
Intrastate branching × Ext. fin. dep.	0.15	0.50
	(0.11)	(0.38)
Covariates	Yes	Yes
R^2 / Log Pseudo Likelihood	35,594	0.30
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:

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

where $\log V\_Total_{ict}$ 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 $\mathbf{Z}_i$ 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

Technical Appendix: Table 5

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

$$(4) \qquad P(N_{st} = n) = \begin{cases} p + (1-p)\left(1+\dfrac{\lambda}{\tau}\right)^{-\tau}, & n = 0 \\ (1-p)\dfrac{\Gamma(n+\tau)}{n!\Gamma(\tau)}\left(1+\dfrac{\lambda}{\tau}\right)^{-\tau}\left(1+\dfrac{\tau}{\lambda}\right)^{-n}, & n = 1,2,\ldots \end{cases}$$

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

$$p = \frac{\exp(\pi_1 \text{Interstate\_Bank}_{st} + \pi_2 \text{Intrastate\_Branch}_{st} + \mathbf{X_{st}}\psi + \omega_s + \tau_t)}{1+\exp(\pi_1 \text{Interstate\_Bank}_{st} + \pi_2 \text{Intrastate\_Branch}_{st} + \mathbf{X_{st}}\psi + \omega_s + \tau_t)}.$$ [20]

[20]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, $\mathbf{X}_{\mathbf{st}}$, 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. 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