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

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Abstract:

Starting in the late 1970s, many U.S. states adopted intrastate branching and interstate banking deregulations that lifted restrictions on bank expansions and facilitated access to finance. We evaluate the effects of these deregulations and the accompanying reduction in credit constraints on foreign firms investing in the U.S. We provide direct, micro-level evidence from U.S. inbound foreign direct investment (FDI) transactions showing that facilitating access to local finance increases the entry rate of foreign multinationals and the number of transactions while reducing the average transaction value, i.e. when credit constraints are relaxed, smaller FDI transactions stand a greater chance of being completed. We also find that the impacts are more pronounced for multiple-transaction investors as opposed to single-transaction investors. Finally, we demonstrate that following banking deregulation the entry rate and the value of transactions increase in industries that are more dependent on external finance relative to industries less dependent on external finance.

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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 deregulations that lifted restrictions on bank expansions. These two types of deregulations 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 increases the entry rate of foreign multinationals as well as the number of FDI transactions while reducing the average FDI transaction value, i.e. when credit constraints are relaxed, smaller FDI transactions are more likely to be completed. A larger number of transactions combined with a lower average transaction value leads to a small aggregate FDI effect, which is statistically insignificant from zero, underscoring the importance of using micro-level data in identifying the impact of access to credit on FDI flows.

[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.

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 have a positive impact on FDI activity. We know from the extant literature that multinational firms utilize significant amounts of 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 deals are more likely to be financed with cash as opposed to stock, and cash transactions in turn are likely to involve external borrowing. Beyond the initial transaction, debt is also extensively used to finance the continued operations of foreign affiliates, which typically use a mix of internal and external local debt financing. Common motivations for the use of local finance include currency risk hedging, transfer pricing and availability of cheap credit. 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. 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. Using a comprehensive dataset of all foreign affiliates of U.S. multinationals, Desai *et al.* (2004) report that foreign affiliates have an external borrowing to assets ratio of over 44 percent. Furthermore, 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. 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'.

To illuminate the mechanism for the effect of banking deregulation, 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 U.S. collateral, which first time investors lack. Further, 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 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 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.[3] The data also provide details on the type of transaction – e.g. new plant, merger and acquisition, or joint venture. We restrict

[3] 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).

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 credit constraints from potential confounding factors. Because of the richness of the data, we are also able to control for a number of transaction- and investor-specific characteristics that may affect the average transaction value, such as the nationality of the foreign investor and the type of transaction. Our econometric models additionally include a host of state-level, time-varying covariates, such as the gross state product (and its growth rate), the unemployment rate, population density, the corporate tax rate, the average wage, the number of foreign trade zones, and market potential, all of which may affect FDI activity and be correlated with banking deregulation. Our results are robust to the inclusion of state-specific trends that additionally allow FDI trajectories to differ across states, as well as country-specific time effects and a host of variables characterizing investor experience. A major advantage of our study compared with cross-country studies is that we are implicitly able to control for many characteristics common to all states, such as macroeconomic policy and federal legislation (with respect to labor and capital markets as well as trade policy) that can affect FDI.

Our findings show that whereas the deregulation of intrastate branching does not have any significant effect on FDI inflows along the extensive or the intensive margin[4], interstate banking deregulation is associated with a higher entry rate of foreign multinationals, a larger number of FDI transactions, and a smaller average transaction value. In particular, 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. When we evaluate the impact of interstate banking deregulation on the size distribution of foreign transactions, we find that above-median value transactions experience greater declines. Our results indicate that with improved access to external finance, foreign firms are able to undertake projects of smaller value.

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. We further study the differential impact of the two banking deregulations in industries that are *more* versus *less* dependent on external finance and find that while transaction values in sectors that are more dependent on external finance are buoyed by the interstate banking deregulation, transaction values in sectors less dependent on external finance fall, i.e. when access to local finance improves, transaction values in sectors *more* dependent on external finance increase compared to transaction values in *less* dependent sectors.

Investigating the impact of banking deregulation along the extensive margin, we find that

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

states that adopted the interstate banking deregulation have experienced a 17 percent increase in the number of inbound FDI transactions, translating to 1.18 new FDI transactions per year. We find that interstate banking deregulation is associated with an increase in the entry rate of foreign multinationals from 1.15 to 1.63 percent. Our results suggest that both entry rates and the number of new inbound FDI transactions increase with the adoption of interstate banking deregulation. We find that both single- and multiple-transaction investors experience an increase in the number of completed transactions and entry rates, but the effect is significantly larger in magnitude for the multiple-transaction investors. Turning to the effect of banking deregulation on FDI activity in sectors that are *more* dependent on external finance, we find that following interstate banking deregulation, entry rates rise by more in sectors that rely *more* heavily on external finance.

Taking the intensive and extensive margin effects together, our results imply that by improving access to local finance, banking deregulations enabled more foreign firms to invest in the U.S., and allowed them to undertake smaller projects on average. Our results also have aggregate-level implications. We show that the larger number of transactions combined with a lower average transaction value lead to an aggregate effect, which is small and statistically insignificant. This finding highlights the importance of using micro-level data in this context.

While we find that interstate banking deregulation has a positive effect on the entry rate and the number of cross-border transactions and leads to smaller transaction values, our analysis suggests that the intrastate bank branching deregulation has 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 margins 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 deregulations. 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 increase and smaller value transactions become more prevalent, however the magnitudes of the effects are greater possibly because multinational firms have a wider array of investment location choices.

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

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

[5] 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).

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 deregulations.

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. The regression results show that there is not an economically or statistically significant relationship between initial FDI presence and the timing of the adoption of banking deregulations across states.

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

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.[6] 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.[7] Each transaction is also classified into one of six modes of entry:

[6] After 1994, only the Bureau of Economic Analysis (BEA) collects such data, however their data are confidential and not publicly available.
[7] 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.

merger and acquisition, new plant, plant expansion, equity increase, joint venture, and other (see Table 2).

To analyze the impact of the two banking deregulations on the intensive margin of FDI, we employ the data on transaction values. 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 in each state-year cell. 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).

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. 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. 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 it likely has little effect on the results. First, assuming selection on observables, we use inverse probability weighting to demonstrate that the results along the intensive margin remain largely unchanged. Second, 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.[8] The estimated impacts of the banking deregulation reforms across the two datasets are very similar,

[8] 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.

further suggesting that omitting transactions with missing values may not bias the estimates much.

The next section provides details on our econometric strategy and describes the different state-level time-varying covariates that may affect either the FDI transaction value or the entry rate of foreign multinationals (as well as the number of new inbound FDI transactions). 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.

Summary statistics for all variables included in our analysis are presented in Panels A and B of Table 2. 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. On average, there are about 6.93 new FDI projects annually (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.

4. Econometric Strategy

4.1 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:

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

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 two indicator variables $\text{Intrastate\_Branch}_{st}$ and $\text{Interstate\_Bank}_{st}$ equal to one starting in the year in which the 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 corporate tax rate and (6) the number of state 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).[9]

Vector $\mathbf{Z}_i$ includes four investor and transaction-specific covariates. First, we allow for 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

[9] Haung (2008) implements an alternative estimator that relies on the geographic discontinuity of intrastate banking deregulations. He compares economic performance of two 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 do not have county-level geographic information (almost all transactions do have information on the state). Second, most FDI transactions likely 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).

transactions may indicate a high-value investor, so one would expect that higher past averages will 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, we include state fixed effects ω_s, in order to control for unobservable, time-invariant, state-specific characteristics that affect the value of inward FDI transactions and may be correlated with the bank branching deregulations. In addition, we include year fixed effects τ_t, to capture economy-wide shocks that affect all states. 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, to allow for cross-state differences in trends of FDI transaction values, 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 and this could affect the value of local firms acquired by 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). The last term in our regression equation (1), ε_{imcstj}, denotes the residual. 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).[10] Note that these weights are time invariant and hence are not affected by the banking deregulations over time. The weights are used so that we can 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.

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

[10] 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 7).

estimator of the variance-covariance matrix is consistent in the presence of any correlation pattern within states over time.

4.2 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 fraction 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. It 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:

$$(2) \qquad \text{Entry\_Rate}^*_{\text{st}} = \alpha_1 \text{Interstate\_Bank}_{st} + \alpha_2 \text{Intrastate\_Branch}_{st} + \mathbf{X}_{\mathbf{st}}\mu + \omega_s + \tau_t + \omega_s * Trend_t + \kappa_{st}$$

$$(3) \qquad \text{Entry\_Rate}_{\text{st}} = \max\{0, \text{Entry\_Rate}^*_{\text{st}}\},$$

where Entry_Rate^*_{st} is the underlying latent variable, which is not observed, and it satisfies the classical linear model assumptions and Entry_Rate_{st} is the observed outcome, defined as the number of new inbound FDI transactions in state *s* and in year *t* divided by the total number of foreign multinationals operating in that state and year (see the Data section and footnote 7 above). Equations (2) and (3) above imply that the observed variable, Entry_Rate_{st}, equals Entry_Rate^*_{st} when $\text{Entry\_Rate}^*_{st} \geq 0$, and $\text{Entry\_Rate}_{st} = 0$ when $\text{Entry\_Rate}^*_{st} < 0$. The vector of state-specific, time-varying controls, $\mathbf{X}_{st}$, is the same as the one used in the model for the transaction values. As before, we include a full set of state and year dummies. We estimate the Tobit model using maximum likelihood. The standard errors are adjusted for heteroskedasticity and are clustered by state. As we do in the case of the intensive margin, we weight by the natural logarithm of the state average manufacturing employment in foreign multinationals over the period 1977-1985. Qualitatively and quantitatively similar results are obtained in unweighted regressions.

In addition to estimating the impact of the two banking deregulations on the entry rate of 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 below 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 FDI transactions. Note that while analyzing the number of new transactions in addition to 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 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.

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

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

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 $\tau \rightarrow \infty$, and approaches the negative binomial as $p \rightarrow 0$. 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

$$p = \frac{\exp(\pi_1 \text{Interstate\_Bank}_{st} + \pi_2 \text{Intrastate\_Branch}_{st} + \mathbf{X_{st}}\psi + \omega_s + \tau_t)}{1 + \exp(\pi_1 \text{Interstate\_Bank}_{st} + \pi_2 \text{Intrastate\_Branch}_{st} + \mathbf{X_{st}}\psi + \omega_s + \tau_t)}.$$ [11]

The vector of state-specific, time-varying controls, $\mathbf{X_{st}}$, is the same as the one used in the model for the 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.

5. Results

This section presents our empirical results. We first consider the impact of the two banking deregulations on the average value of inbound FDI transactions, as well as on the entire distribution of values. We then turn to 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. Finally, we discuss the aggregate implications of the banking deregulations, by evaluating their impact on the aggregate volume of inbound FDI.

[11] 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.

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 transaction value and the entry rate, 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:

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

The dummy variables take on a value of one in the i^{th} year before or after the deregulation and are zero otherwise.[12] 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 (log of) transaction values or entry rates 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, one can clearly see a declining average transaction value and increasing entry rate following the interstate banking deregulation.

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

5.1 Impact on the Average FDI Transaction Value

Table 3 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 3 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 (= $(e^{-0.35}-1)*100$) percent following the adoption of the interstate banking deregulation.[13] This finding suggests that the increased availability of credit as a result of the interstate banking deregulation allowed smaller FDI transactions to take place – a more efficient and competitive banking system provided financing to smaller projects that otherwise may not have received financing. This, in turn, lowered the average value of foreign direct investment transactions in the U.S. This 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. Our finding also complements previous work on Japanese 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 marginally positive but insignificant.[14] The lack of a significant effect from the

[13] 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 $(e^{-0.35}-1) = -0.295$. Note that for estimated coefficients that are small in magnitude, this procedure makes little difference.

[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

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 this 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 is also most likely true for multinational companies investing abroad. Furthermore, national banks may have a comparative advantage relative to single-state banks in evaluating foreign investment projects.

In the second column of Table 3, 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 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 decreased 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

problem, we re-estimated our empirical specification with the deregulation indicators included one at a time. Compared to the results reported in column (1) of Table 3, there is almost no change in the estimated coefficients on intrastate branching and interstate banking when the indicator variables are included one at a time. The coefficient on intrastate branching remains marginally positive and insignificant, while the coefficient on interstate banking remains negative and significant.

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 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 3, 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 2 for the list of countries).[15] 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'

[15] 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.

investment decisions. Focusing on this exhaustive specification, 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 3, 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 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 (1). The corresponding results are presented in Appendix Table A1. The estimated effects of both banking deregulations are very similar to those reported in Table 3, where we do not correct for missing values.

In Table 4, 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 (1), which corresponds to column (8) of Table 3, on these two subsamples. Columns (1) and (2) of Table 4 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 for the full sample).

Turning to the investor- and transaction-specific covariates in column (3) of Table 4, 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 3. 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.

In columns (4) and (5) of Table 4, 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) and column (5) to column (3) of Table 4.

We next check whether the two banking deregulations have similar effects on transactions of different sizes. To that end, we estimate the most general specification including covariates, state trends, a full set of fixed effects and source country year-effects separately for the transactions that are above and below the median transaction value. The results of this subsample analysis are presented in Panel A of Table 5. They show that the interstate banking deregulation mainly affected larger transactions with values above the median. While the coefficient on interstate banking is negative and significant for the subsample with values above the median, it is small and not statistically distinguishable from zero for the subsample with values below the median.

Secondly, we check how the two banking deregulations affect not only the average transaction value, but also the 25^{th}, 50^{th}, and the 75^{th} percentiles of the value distribution. Because small projects are more likely to benefit from access to credit, we expect the right tail of the distribution of transaction values to shift to the left significantly more than the left tail, i.e., we anticipate that transaction values in the 75^{th} percentile would decline significantly more than values in the 25th percentile as a result of the interstate banking deregulation. We confirm this hypothesis by using quintile regressions to estimate the effect of the two banking deregulations on the three different percentiles of the transaction value distribution. We implement the most general specification that includes covariates, state trends, the full set of fixed effects and the source country year effects described above. The results are presented in Panel B of Table 5. As anticipated, we find that the interstate banking deregulation had larger negative impacts on higher percentiles of the distribution. The coefficients we obtain for the 75th and the 50th percentiles are identical (-0.41) and they are significant at the 1 and 10 percent levels, respectively. The estimate obtained for the 25th percentile is marginally positive and it is not statistically significant. Taken together, the results in Table 5 suggest that the negative impact of the interstate banking deregulation along the intensive margin, i.e., in terms of transaction values, was mainly driven by declines in the values of large transactions.

Finally, we provide even more 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 which are less reliant on external finance. We categorize all 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

[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.

finance investment. We estimate our augmented specification separately for these two groups of transactions – those in industries with negative external finance dependence (*less* dependent) and those in industries with positive external finance dependence (*more* dependent). The results are presented in columns (1) and (2) of Table 8.

The coefficient on interstate banking obtained with the *less* external finance dependent subsample in column (1) is negative and both economically and statistically significant. However, the estimate is marginally negative and insignificant for the *more* external finance dependent subsample. These results imply that the decline in the average foreign transaction value in the overall sample is mainly driven by a larger decline in transaction values in *less* external finance dependent sectors. The average transaction value for this group likely decreased as better access to local credit provided a bigger incentive for smaller foreign firms (compared to larger firms with more free cash flow) in *less* external finance dependent industries to invest in the states that deregulated their banking systems. On the other hand, interstate banking deregulation facilitated both small and large transactions in industries that are heavily dependent on external finance, which led to a small and insignificant change in the average transaction value for those industries.

More importantly, to formally test if the effect of banking deregulation on transaction values changes with the need for external finance, we estimate equation (1) using the entire sample, but we include interaction terms between the deregulation indicators and the continuous external finance measure of Cetorelli and Strahan (1998).[17] Column (3) of Table 8 presents these

[17] The subsample analysis in columns (1) and (2) of Table 8 also shows that the covariates included in the specification have different effects on transaction values in the *less* versus *more* external finance dependent subsamples. To allow for these different dynamics, when we estimate the specification with the external finance interaction terms in column (3), we additionally include interaction terms of the covariates with an external finance dummy that takes on a value one when the transaction belongs to a *more* external finance dependent industry.

results. The main effect of interstate banking remains very similar to the findings in Table 3, 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* 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.

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

In Table 6, we presents the results from the Tobit specification (see equations (2) and (3)) 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. Consistent with our findings from the intensive margin analysis (see the transaction value results in Table 3), the coefficient on the interstate banking indicator is both statistically and economically significant, but the one on interstate branching is not. The estimated coefficient of 0.48 (with a standard error of 0.18) on interstate banking 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 2, 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.

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 which deregulated interstate banking experienced a 0.38 percentage point (or about 33 percent) increase in the 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 6 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 not 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 percent increase is larger since the average entry rates computed when we use only transactions, for which the value is not missing are about twice as low. In general, the banking deregulation impact on the entry rate of foreign multinationals reported in Table 6 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 2) 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 banking deregulations on the number of new inbound FDI transactions. Table 7 presents the results from the zero-inflated negative binomial model (see equation (4)) that estimates the impact of interstate banking and intrastate branching deregulations on the number of FDI transactions. We only report the 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 7. 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 intensive margin (see the transaction value results in Table 3) and the entry rate (see Table 6), the coefficient on the interstate banking indicator is both statistically and economically significant, but the one on interstate 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 ($=(e^{0.30}-1)*100$) following the adoption of the interstate banking deregulation. Since the average number of new transactions (including those with missing transaction values) is 6.93

(see Table 2, 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 7 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). Combined with the intensive margin results for the multiple-transaction investors in Table 4, columns (2) - (5), these findings suggest that the interstate banking deregulation led such investors to complete greater number of transactions with smaller values. The result is consistent with our hypothesis that multiple-transaction investors, which are more likely to have U.S. collateral compared with single-transaction investors, are more prone to be affected by changes in access to local bank finance.

Lastly, we consider the impact of interstate banking and branching deregulations on the entry rate of foreign multinationals in *less* external finance dependent industries compared to their impact on the entry rate in *more* external finance dependent industries. We would expect the impact of banking deregulations to be more pronounced in sectors that are more reliant on external finance (Rajan & Zingales 1998). As in the case for the transaction values we discussed in the previous subsection, we categorize the number of new FDI transactions into two groups based on the two-digit SIC industry-specific external finance dependence measure in Cetorelli and Strahan (2006). We estimate our preferred Tobit specification with covariates and state trends for the entry rates separately for the two groups. Columns (4) and (5) of Table 8 present the results. The estimates show that interstate banking deregulation is associated with higher entry rates in both types of industries. The coefficients on the interstate banking indicator imply that the states that adopted interstate banking experienced an increase in the entry rate of foreign multinationals of 0.27 and 0.13 percentage points in the *more* and *less* external finance dependent industries, respectively, with only the former effect being statistically significant at the 5 percent level. As expected, the effect is larger for *more* external finance dependent industries since relaxing credit constraints provides a greater investment incentive for foreign multinationals operating in such industries.

We formally test whether the impact of banking deregulation varies with the degree of external finance dependence by combining the data on both types of industries (*less* and *more* external finance dependent) and including interaction terms between the external finance dependence dummy and the interstate banking and intrastate branching indicators in the augmented specification.[18] In column (6) 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*

[18] 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.

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, and its interaction with the external finance dummy is positive and highly significant. These results confirm our previous findings and provide further direct evidence of the importance of the local finance channel for foreign investment in the U.S. The estimates suggest that 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 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 it is not significant. Its interaction with the external finance indicator is positive, but also not significant.

5.3 Impact on Overall FDI Volume

Next, we evaluate the effect of banking deregulation on the total volume of inbound FDI across U.S. states. Because we find that the average transaction value declines while the number of transactions grows as a result of the interstate banking deregulation, the effect on the total volume of new FDI inflow is ambiguous. To estimate this impact, we first aggregate the individual transaction values in each state and year into an aggregate total FDI inflow measure. This aggregate number reflects both the number of new foreign investment transactions (the extensive margin) and their values (the intensive margin). Panel A of Table 9 reports the impact of the two banking deregulations on the aggregate FDI inflow. This specification is similar to equation (1), but the dependent variable, aggregate FDI inflow, is not logged because it is zero in a number of instances. The results reveal a positive but statistically insignificant impact of the interstate banking deregulation – aggregate FDI increases by about 120 million 1983 U.S. dollars as a result of the deregulation. When we consider the specification with covariates and state trends in column (3), we find that following the adoption of the interstate banking deregulation, states experienced a slight decline in aggregate FDI inflow of about 20 million 1983 U.S. dollars; however, this effect is statistically indistinguishable from zero. As before, the impact of intrastate branching on total FDI inflows is small and insignificant.

By aggregating the transaction-level data on FDI project values, we can only estimate the impact of banking deregulation on the total value of *new* foreign inbound transactions, i.e., *new* entries. To evaluate the effect of improved access to local financing on the aggregate value of FDI while accounting for both entry and exit of foreign multinationals in the U.S., we estimate our specification using annual state-level FDI stock data (for the sample period 1977-1994) provided by the Bureau of Economic Analysis. We use the change in the total value of plant, property, and equipment of foreign companies operating in the U.S. as our dependent variable.[19] These estimates are reported in Panel B of Table 9.

The first column in Panel B presents the results of the baseline specification without covariates. The coefficient on interstate banking is significant at the 5 percent level, but the coefficient on interstate branching is not. The coefficient of 85.70 implies that interstate banking deregulation is associated with an increase in aggregate net FDI flow of about 86 million 1983 U.S. dollars.[20] However, once we include state-specific covariates, interstate banking becomes

[19] Computing the change in the dependent variable involves losing one year of the panel in the estimation. The change in the total value of plant, property, and equipment gives us the net flow of aggregate FDI in a state for a given year. We choose to use this net flow variable as our dependent variable to facilitate the comparison of the results to the ones we obtain using the new FDI inflow data reported in Panel A of Table 9. Qualitatively we obtain similar results if we use the level or the growth rate of the total value of plant, property, and equipment, instead.

[20] The mean value of the change (year-to-year) in the total value of a state's foreign plant, property and equipment in our sample is 200 million 1983 U.S. dollars.

insignificant, and the size of the coefficient drops to 25.86. It declines further to -0.44 when state-specific trends are included. This corresponds to a very small decline in net aggregate FDI flows of about half a million 1983 U.S. dollars, but this estimate is not statistically significant.

In short, using aggregate FDI data, we fail to identify a significant effect of improved access to local finance on FDI inflows. However, using transaction-level data, we demonstrate that interstate banking deregulation leads to a greater number of transactions, a higher entry rate of foreign multinationals, and a lower average transaction value. When the two effects along the extensive and intensive margins are combined, the overall effect is practically zero, although it is imprecisely estimated. Our findings underscore the importance of using micro-level data in evaluating the impact of access to external finance on FDI flows into the U.S.

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 credit constraints in two important ways. First, we provide direct micro-level evidence from the U.S. that credit constraints affect both the *intensive* and the *extensive* margins of inbound FDI. 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 *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 reduced credit constraints. More specifically, we estimate a difference-in-differences model with multiple time periods and we 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. Specifically, 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 is robust to including a comprehensive list of state-level, time-varying controls and trends, investor- and transaction-specific covariates, as well as source country and mode of entry fixed effects. Further, when we evaluate the impact of the interstate banking deregulation on the size distribution of foreign transactions, we find evidence that the largest reductions occurred in the higher percentiles of the distribution. These results suggest that when access to local credit improves as a result of increased bank competition, foreign firms are able to undertake smaller inbound FDI projects.

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. Taken together, these results imply that by facilitating access to local external finance, U.S. banking deregulation has enabled more foreign firms to invest in the U.S. and allowed them to undertake smaller projects on average.

Finally, to shed light on the mechanism for the effect of 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 U.S. collateral, 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 banking deregulation should be more pronounced in industries that are more reliant on external finance, which is what we find. Compared to projects completed in industries less reliant on external finance, both transaction values and entry rates for projects completed in industries that are *more* dependent on external finance increase after the interstate branching deregulation was adopted.

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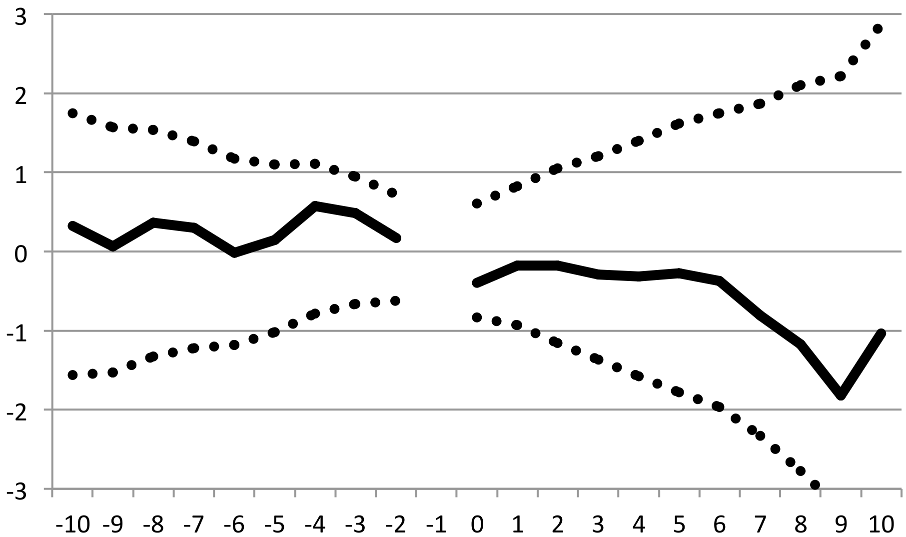
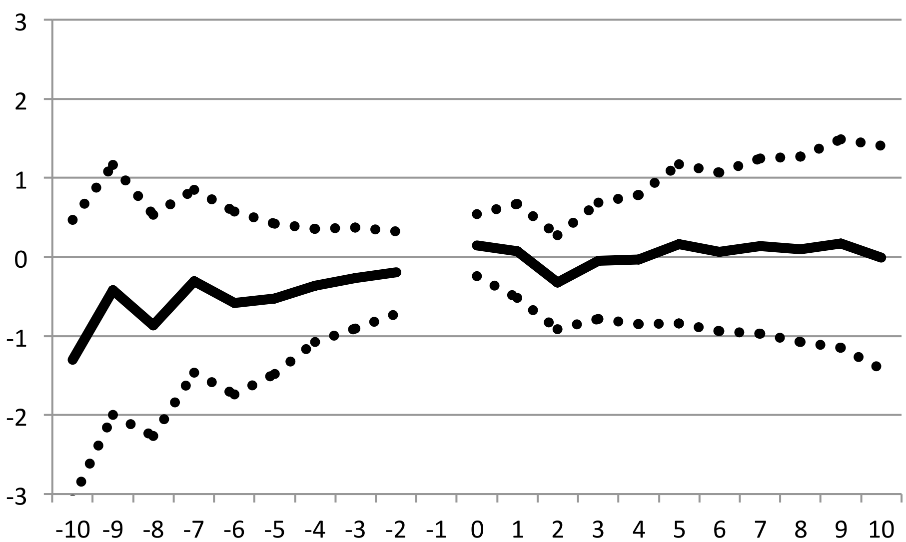
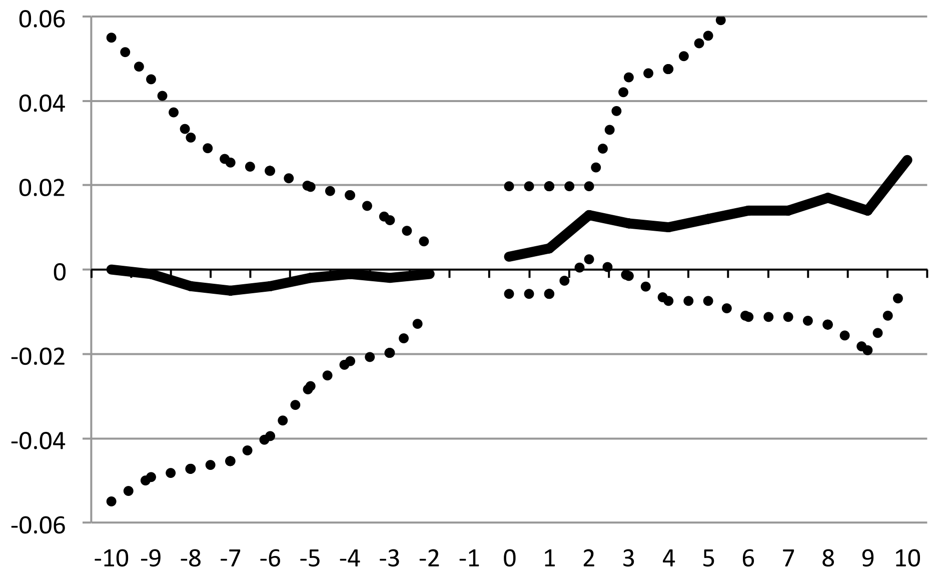


Interstate Banking

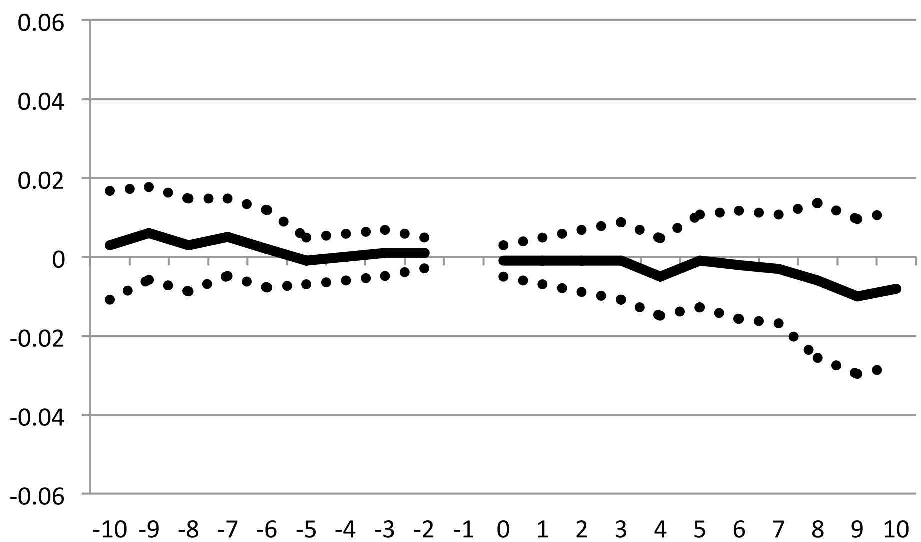


Intrastate Branching

Figure 1. 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: $\pm 0.5 + \frac{95\%\ \text{Confidence Interval}}{3.5}$.



Interstate Banking



Intrastate Branching

Figure 2. 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: $\frac{\text{95\% Confidence Interval}}{0.5}$.

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. 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 2. 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 3. 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:

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

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 $\boldsymbol{Z}_i$ 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 4. 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 5. The Impact of Interstate Banking and Branching Deregulation on the Distribution of Foreign Direct Investment Transactions in the U.S. Manufacturing Sector, 1977-1994.

Panel A: The Impact of Banking and Branching Deregulations on Large (above Median Transaction Value) and Small (below Median Transaction Value) transactions. We split the sample into transactions that are above and below the median transaction value. We estimate the most general specification including covariates, state trends, a full set of fixed effects and source country time effects separately for the two subsamples.

Dependent Variable – *ln*(Transaction Value)

Variable	(1)	(2)
	Values above the median	Values below the median
Interstate banking	-0.32***	0.08
	(0.12)	(0.09)
Intrastate branching	0.16	0.06
	(0.19)	(0.12)
R^2	0.27	0.25
No. Obs.	1,457	1,458

Panel B: Quintile Regressions. We use quintile regression to estimate the effect of banking deregulation on the, 25th, 50th, and the 75th percentiles of the transaction values distribution. We estimate the most general specification that includes covariates, state trends, a full set of fixed effects and source country time effects. Block-bootstrapped (by state) 95-percent confidence intervals reported in brackets. *** 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)
	25th percentile	50th percentile	75th percentile
Interstate banking	0.06	-0.41*	-0.41***
	[-0.69 0.44]	[-0.79 0.05]	[-0.99 -0.06]
Intrastate branching	0.11	0.05	0.19
	[-0.33 -0.65]	[-0.34 0.46]	[-0.35 0.91]
Pseudo R^2	0.19	0.19	0.21
No. Obs.	2,915	2,915	2,915

Table 6. 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. 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 7. 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 in each state (including those with missing values). 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 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). Columns (1) through (3) use the natural logarithm of the transaction value as dependent variable and all three specifications include source country, mode of entry, and industry dummies, as well as source-country-specific time effects. Results of the augmented specification for the two subsamples of transactions are presented in columns (1) and (2). Column (3) presents results from a specification using the full sample and includes interaction terms between the deregulation indicators and the continuous external finance measure. The specification in column (3) also includes interactions of the covariates with the external finance dummy that takes on a value of one when the transaction belongs to a more external finance dependent industry. Columns (4) and (5) present Tobit results for transactions in industries more and less dependent on external finance, respectively, using the entry rate of foreign multinationals (new FDI transactions/number of foreign multinationals) as dependent variable and normalizing the two indicator variables Interstate banking and Intrastate branching by 100. Column (6) uses the full sample and includes interaction terms between the external finance dependence dummy and the interstate banking and intrastate branching indicators. 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. All specifications include state and year dummies, as well as state trends. 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)	(3)	(4)	(5)	(6)
	ln(Transaction Value)			Entry Rate		
	Less Ext. Fin. Dep.	*More* Ext. Fin. Dep	All	*Less* Ext. Fin. Dep.	*More* Ext. Fin. Dep.	All
Interstate banking	-0.51**	-0.07	-0.28***	0.13	0.27**	0.03
	(0.22)	(0.17)	(0.09)	(0.11)	(0.13)	(0.12)
Interstate banking × Ext. fin. dep.			0.96**			0.344***
			(0.46)			(0.13)
Intrastate branching	-0.13	0.22	0.10	-0.06	0.06	-0.06
	(0.29)	(0.18)	(0.17)	(0.09)	(0.16)	(0.109)
Intrastate branching × Ext. fin. dep.			0.50			0.15
			(0.38)			(0.11)
Covariates	Yes	Yes	Yes	Yes	Yes	Yes
R^2 / Log Pseudo Likelihood	0.47	0.29	0.30	16,826	19,736	35,594
No. Obs.	909	2,006	2,915	828	828	1,656
Zero Obs.	-	-	-	346	247	593

Table 9. The Impact of Interstate Banking and Branching Deregulation on Aggregate FDI Flows in the U.S. Manufacturing Sector.

Panel A. Aggregate New FDI Inflows, 1977-1994.
We aggregate the individual transaction values in each state and year into an aggregate total FDI inflow measure. Column (1) presents the results from an OLS specification that includes state and year fixed effects. The augmented specification in column (2) includes covariates, and column (3) further includes state-specific time trends. 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 – Aggregate New FDI Inflow			
Variable	(1)	(2)	(3)
Interstate banking	119.89	27.39	-19.87
	(99.84)	(120.60)	(124.19)
Intrastate branching	-75.47	-69.28	-29.67
	(89.41)	(68.98)	(86.85)
Covariates	No	Yes	Yes
State trends	No	No	Yes
R^2	0.35	0.39	0.42
No. Obs.	828	828	828

Panel B. Change in Aggregate Plant, Property, and Equipment (PPE) in Foreign Multinationals (Net Aggregate FDI Flow), 1978-1994.
We construct a measure of net aggregate FDI flows that account for entry and exit of foreign multinationals from Bureau of Economic Analysis data. We use the change in the total value of plant, property and equipment (PPE) of foreign companies operating in the U.S. as our dependent variable. Column (1) presents the results from an OLS specification that includes state and year fixed effects. The augmented specification in column (2) includes covariates, and column (3) further includes state-specific time trends. 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 – Change in the PPE of foreign companies operating in the U.S.			
Variable	(1)	(2)	(3)
Interstate banking	85.70**	25.86	-0.44
	(34.14)	(44.69)	(43.56)
Interstate branching	13.65	19.15	8.86
	(44.87)	(43.30)	(40.431)
Covariates	No	Yes	Yes
State trends	No	No	Yes
R^2	0.36	0.40	0.44
No. Obs.	782	782	782

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 is 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