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Abstract:

National borders are still strong barriers for mergers and acquisitions in Europe. We estimate a gravity equation model based on NUTS 2-regions and find that the restraining impact of national borders decreased by about a third between 1990 and 2007. However, there has been no significant change since 1997, i.e., two years before the introduction of the Euro. To benchmark our results we run a corresponding analysis within the United States using the ten federal OMB regions as country equivalents. The ‘quasi border’-effect in the US is weaker than in the EU and even declines more during the same time period. We conclude that European integration policy has little effect on fostering cross-border transactions.

JEL-Classification: F21, G34

Key Words: European integration, corporate control, border effects

Disclaimer: The paper presents the personal opinion of the author and does not necessarily reflect the views of the Deutsche Bundesbank.

I. Introduction

Over the last decades the integration of European capital markets has been one of the top priorities for policy makers. In January 2000, a report by the European Commission on product and capital markets, also known as the Cardiff report, was pleased to announce a strong increase of cross-border transactions within the EU in previous years. The commission stated that 'cross-border investment and mergers and acquisition activity […] are the most dynamic factors driving integration in today's Internal Market' (see COM(2000) 26, p.9). Five years later, cross-border investments were further encouraged by integration politics with 'a view to the completion and functioning of the single market', as manifested in the EU cross-border directive 2005/56/EC (p.1). This empirical study focuses on European integration in the market for corporate control, which is economically important for the following reasons. First, an integrated market would foster acquirers' search for the best suited target firm with the largest possible synergies. Arguably, resulting business combinations should be more efficient in an integrated European market than in segmented markets. The EU Commission also cites better dissemination of knowledge and technology as a benefit (Ilzkovitz et al. 2007). Second, both the threat of hostile takeovers and actual transactions help to discipline management. An integrated European market for corporate control could thus mitigate potential corporate governance shortcomings (Rossi and Volpin 2004). The lesser the market integration, the fewer potential benefits can be reaped.

We analyze European M&A transactions from 1990 to 2007 and find a strongly negative and highly significant border effect. We include all transactions in the 15 countries that formed the EU in 1995 and construct a gravity equation model explaining M&A investment flows between European regions. This model allows us to estimate the border effect, while controlling for various factors like distance, local economic activity, cultural and legal differences, and country and time effects. To track the border impact over time we introduce a panel analysis and find an overall decreasing restraining-effect imposed by national borders. It is not obvious, however, whether the declining border effect can be attributed to European integration politics or, e.g., general technological trends. We compare the European results with developments within the US by using the ten Federal Regions defined by the Office of Management and Budget as country-equivalents in the US similar to countries in Europe. Not surprisingly, we find M&A activity in the US far more integrated than in

Europe. By running the same panel analysis on US data we find an even stronger integration trend over time within the US, i.e., a lessening impact of the—artificially constructed—borders. While the border effect, measured by annual coefficients in the gravity model, decreases by 33.5 percent in Europe, the 'border effect' within the US decreases by 45.8 percent over the same time period. Our findings indicate that the European market still is much more fragmented in contrast to the US and integrated at a slower speed than the US market over the last fifteen years. This places some doubt on the impact of political measures for European integration in the market for corporate control. Our results remain robust when we introduce stock market performance or spatial dependencies as additional control variables and when we distinguish between small and large companies. Also, our findings do not change when we use the number of investments per region instead of the total volume as the dependent variable and when we estimate negative binomial regression models instead of Poisson models.

Our paper rests upon two strands in the literature, namely the analysis of foreign direct investments' locations, and the exploration of borders' impact on economic activity. The latter research started with McCallum's (1995) analysis of the US-Canadian border's impact on trade between regions. Engel and Rogers (1996, 2001) estimate the borders' effects on relative prices and find a strong impact. Chen (2004) has examined European borders' influence on trade within the EU and finds considerable restraining effects induced by borders. Balta and Delgado (2008) report that the home bias in trade within the EU has barely changed in recent years. There is, however, evidence that a home bias in trade also exists on a regional level within the US (Wolf 2000), i.e., borders might not be solely responsible for excessive intra-national trade. The second pillar our paper rests on is the large body of literature that is concerned with the features that help in attracting foreign direct investments or M&A transactions. Geographical characteristics such as distance, adjacency, and time zones as well as culturally grounded variables such as a common language, similar legal systems, and other cultural features have been found to influence investments from one country to another (see Stein and Daude 2007). Huizinga and Voget (2009) show that tax systems do have an influence on the location of headquarters after mergers and the level of M&A activity in general. Rossi and Volpin (2004) find that the level of investor protection positively influences international M&A activity in a country. Studies find that M&A transactions depend on tariffs (see Hijzen et al. 2008), but given the introduction of the European tariff union already in 1968, this is not a relevant topic for our paper. The workhorse model in this literature is the gravity model (see Head and Ries 2008) that has recently been subject to considerable modifications and extensions (see Anderson and van Wincoop 2003; Silva and Tenreyro 2006). We use this model, incorporating those developments which in general make results more robust and lead to more conservative estimations of border effects than the McCallum specification.

In contrast to other studies that analyze foreign direct investments (FDI) in general, we only take brownfield investments into account, i.e., M&A transactions. Thus, we are not able to comment on European integration in case of greenfield FDI. Nevertheless, most FDI is covered by our analysis since in the developed world on average 83 percent of all FDI has taken the form of M&A between 1997 and 2007 (UNCTAD 2009, own calculations). By focusing on M&A transactions and simultaneously taking domestic deals into account, we are able to control for distance effects and approximate border effects within the EU. There is sufficient documentation that the 'home' or 'regional' bias which has been established for equity holdings (see French and Poterba 1991) and trade (Wolf 2000) also influences M&A transactions: firms tend to buy other firms nearby (Kang and Kim 2008; Uysal et al. 2008). Without accounting for this tendency, measures of border effects could be biased. Domestic and cross-border transactions might have different motives, but do not show very different results: As for domestic deals, studies of acquirer returns in cross-border transactions deliver mixed results. Eckbo and Thorburn (2000) find mostly negative but insignificant returns for US acquirers buying Canadian targets, Goergen and Renneboog (2004) find positive acquirer returns in Europe, and Moeller and Schlingemann (2005) a negative cross-border effect for US acquirers. In a recent paper complementary to ours Coeurdacier et al. (2009) find that the introduction of the European Monetary Union increased cross-border M&A towards the Euro-area by 80%, and intra-Euro transactions by 160% between 1985 and 2004 based on a country level analysis. In our analysis, we use smaller regions instead of countries to shed more light on the dynamics of intra-European transactions.

This paper contributes to the existing literature in several ways. To our knowledge, this study is the first to estimate the development of border effects in gravity models over time. Instead of country data, this study uses regional data and thus allows for regional differences in economic characteristics and activity. Finally, it is the first to provide a benchmark for the relative assessment of European border effects.

II. Data and Methodology

We take M&A transaction data from Thomson Reuters' One Banker database. Our sample consists of all mergers and acquisitions with an effective transaction date between January 1990 and December 2007 where both, target and acquiring company, are located within the EU15, the member countries of the European Union prior to the accession of ten candidate countries in May 2004. The EU15 comprised the following 15 countries: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden, and United Kingdom. Even though Austria, Finland and Sweden did not join the EU until 1995, we include all data starting in 1990. Excluding them or restricting our analysis to the period of 1995 until 2007 yields qualitatively unchanged results. Our data starts in 1990 which was the year when the EU merger regulation was officially enacted by Council Regulation (EEC) No. 4064/89, and we end our observation period in December 2007, the transposition deadline for EU member countries to enforce the cross-border directive 2005/56/EC (several countries did not comply at the time).

In our sample, we include transactions by private and publicly listed companies where more than 50 percent of target stock is acquired, and where the locations of target and acquiring companies' headquarters are known. We use the acquirer's ultimate parent company as acquiring company since in many transactions the reported acquiring company is a subsidiary of a corporation. We exclude private equity firms and other professional investment firms (SIC 6700 to 6799) since they have different business models and investment criteria to other firms. We further restrict our sample to transactions with a recorded deal value (in Euro). These criteria result in a sample of 8857 transactions. We do not distinguish between industries since our unit of measure is the aggregated investment flow between regions within the EU. We adjust all currency-denominated values for inflation to year 2000 levels by the annualized Harmonised Index of Consumer Prices from Eurostat for the EU and the annual average Consumer Price Index (CPI) published by the Bureau of Labor Statistics for the US.

We use data on a regional base since recent research has shown a trend for acquirers to buy nearby targets within countries because of better information, lower transaction costs and higher expected synergies (Kang and Kim 2008; Uysal et al. 2008). By neglecting these findings country-level analyses might overstate the impact of borders on M&A transactions. We geocode all address data and assign each company’s location to its respective region based on the European Union’s ‘Nomenclature des unités territoriales statistiques’ (NUTS region classification) on level 2. NUTS levels include the national (NUTS 0), larger regional (NUTS 1), and the NUTS 2 level with a population per region of between 800,000 and 3 million people. We measure distances between regions as centroid-to-centroid distances in kilometers.

Table 1: M&A Activity in Europe 1990-2007

This table shows a matrix of all M&A investments in our sample between 1990 and 2007 with known transaction value. Rows contain target and columns contain acquirer countries, thus domestic transactions form the diagonal. The figures show the heterogeneity among countries’ M&A activities as well as generally skewed distribution towards domestic transactions.

Target nation	Acquirer nation															
	AT	*BE*	*DE*	*DK*	*ES*	*FI*	*FR*	*GR*	*IE*	*IT*	*LU*	*NL*	*PT*	*SE*	*UK*	*Sum*
AT	23	2	9	0	1	3	1	0	1	3	0	4	0	4	4	55
BE	0	49	5	1	4	1	18	1	3	3	0	11	0	5	46	147
DE	18	12	257	6	9	13	36	2	8	23	5	25	1	34	196	645
DK	0	0	4	66	1	4	7	0	4	2	0	3	0	30	27	148
ES	0	3	22	1	334	0	33	0	2	21	2	16	16	9	63	522
FI	2	0	3	3	2	94	2	0	0	1	1	1	0	26	12	147
FR	0	30	49	7	21	7	454	0	7	35	0	21	1	16	196	844
GR	0	0	0	0	0	0	2	16	0	0	0	0	0	0	1	19
IE	0	2	3	1	0	0	0	0	64	1	0	1	0	1	70	143
IT	1	7	19	1	11	2	26	1	0	324	0	10	0	10	60	472
LU	0	0	3	0	0	0	1	1	0	0	2	0	0	0	1	8
NL	2	19	15	7	4	5	12	1	5	8	2	107	0	10	108	305
PT	0	1	2	0	8	0	3	0	0	2	0	2	30	2	3	53
SE	2	2	17	17	2	25	12	0	3	1	0	10	0	254	52	397
UK	3	22	79	14	19	11	109	2	114	22	2	53	0	42	4460	4952
Sum	51	149	487	124	416	165	716	24	211	446	14	264	48	443	5299	*8857*

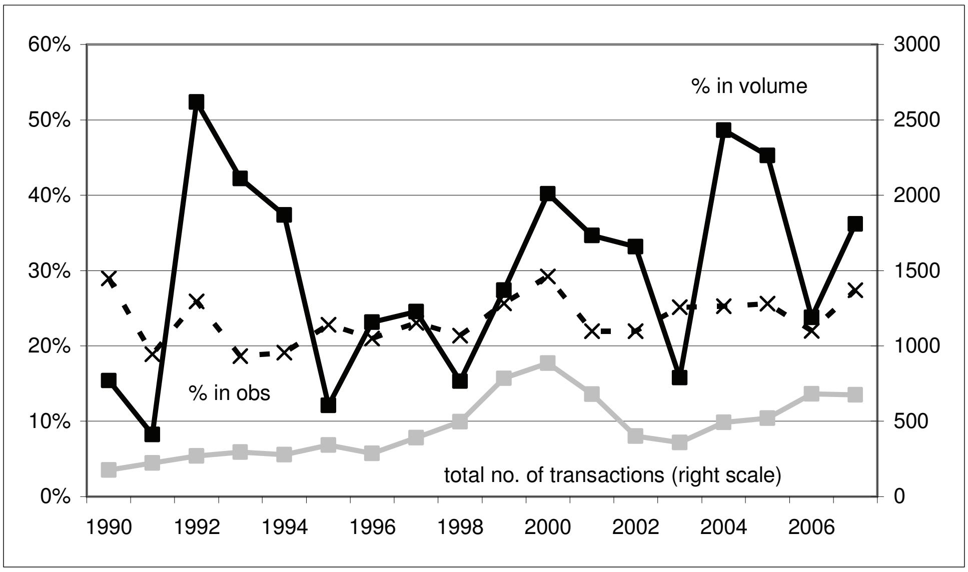
Table 1 shows the transaction in a country matrix with target countries in rows and acquirer countries in columns. In total, transactions are skewed towards domestic transactions and are distributed heterogeneously between countries. Most activity takes place in Great Britain with more than 50 percent of all transactions. Great Britain also has the most cross-border transactions in absolute terms (839 acquisitions in other EU15 countries). Nevertheless, these cross-border transactions comprise only 16 percent of UK's total M&A activity. Austria and Belgium have the most international market for corporate control; both have more international than domestic deals.

Our main focus is on the development of, and potential barriers to, international transactions within the EU. Figure 1 illustrates the total number of M&A transactions by announcement date per year (grey line, right scale). Starting with 176 transactions in 1990 M&A activity hit an all-time high in 2000 with a total number of 887 deals. After a period of considerable lower activity, the number of deals in our sample is up again to 675 in 2007.

The black straight line in Figure 1 marks the development of the share of cross-border transactions measured by transaction volume. In 1990, 15 percent of all Intra-European deals in our sample are cross-border (29 percent in number of transactions). There is a slight upward trend in the share of cross-border deals, though the numbers fluctuate considerably per year. The dotted black line shows cross-border transactions as a percentage of the total number of deals. From 1990 to 2000 the shares of the number of cross-border transactions first decrease and then increase again, followed by a sharp decrease to below 22 percent in 2001. Only in 2007 the share of cross-border M&A deals reaches a level of above 27 percent again. The cross-border percentages measured by the number of deals weight all transactions equally and thus put higher emphasis on smaller companies. In what follows, we stick mostly to the volume figures and use count-based measures only as a robustness check (the results remain qualitatively the same).

Figure 1: Percentage of Cross-Border M&A Activity 1990-2007

Figure 1 shows the development of cross-border M&A transactions over time. The straight black line shows the percentage of cross-border deals in terms of deal volume; the dotted black line the share of cross-border transactions in numbers. The light gray graph depicts the total number of M&A transactions. We base the time series aggregation on the year of announcement and adjust all currency-denominated values for inflation to year 2000 levels by the annualized Harmonised Index of Consumer Prices from Eurostat

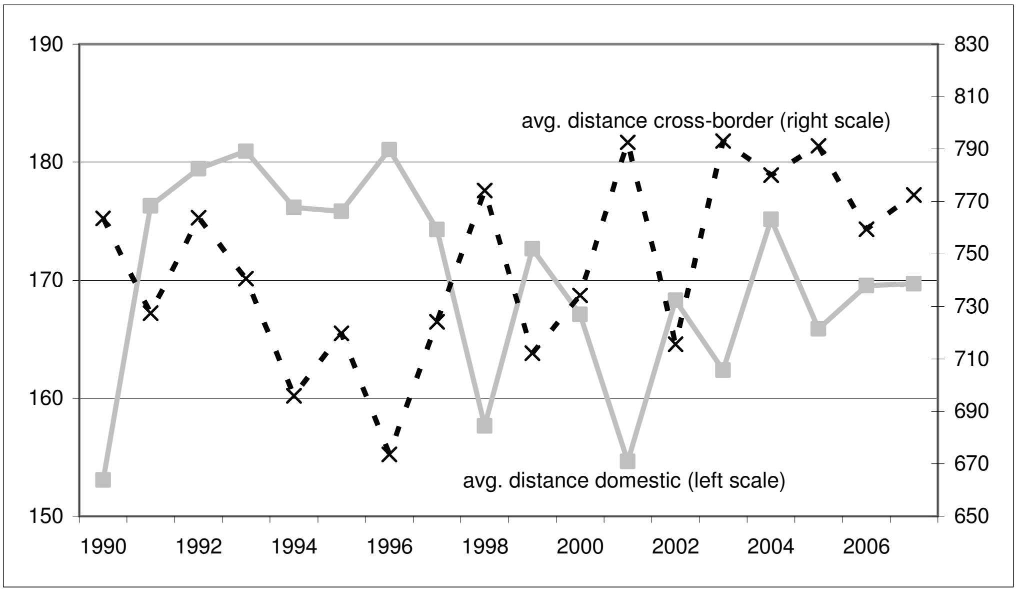


There is no clear-cut trend in terms of the share of cross-border activity. The volume graph shows higher volatility, since a few high-priced transactions can easily distort the mean. The graph shows early high levels in 1992, 1993 and 1994 (above 37 percent). We have some doubts, however, concerning the data quality of Thomson Financial's data base regarding continental European deals in the early 1990s. Thus, we do not want to overemphasize the peak in the beginning of this period. The mid nineties saw low levels of cross-border activity (below 15 percent) and a small peak of 40 percent in the year 2000. After 2000, there is a decline until 2003 and a strong increase in the percentage of cross-border transactions up to almost 50 percent in 2004 and 2005.

Transaction distances form the core of gravity models. Figure 2 shows the average headquarter-to-headquarter distance between the target and the acquirer's ultimate parent company in kilometers. We separate cross-border deals (dotted black line, right scale) and domestic deals (grey line, left scale). For instance in 1991, acquirers purchased companies in other European countries that were located on average 727 kilometers away, while domestic transaction partners were on average 176 kilometer apart. On average, there is not a clear-cut trend in the distances between headquarters in European M&A. Domestic deals' distances decrease slightly where cross-border distances show an increasing trend (median distances give roughly the same picture).

Figure 2: Average Transaction Distance

In this figure we show the average transaction distance as measured by the headquarters' distance between target and acquirer ultimate parent companies in kilometers. The dotted black line shows the yearly average headquarter-to-headquarter distances of cross-border transactions (right scale). The grey straight line shows distances for domestic deals (left scale).



Cross-border M&A activity in Europe fluctuates considerably over time, measured both by the share of cross-border deals and the distance between transaction partners. However, between the mid-nineties and the end of our observation period there is some increase in the share of cross border deals as well as in the distance between transaction partners.

For a deeper understanding of European cross-border investments we turn to the gravity equation approach and aggregate all transaction data on NUTS 2 level. In the M&A and FDI literature countries are often chosen as the unit of measure. In contrast, the NUTS 2 level allows for a detailed analysis of the distance between two regions and local variations in economic activity. As stated before, we use the sum of transaction values in Euro to measure investment flows between two regions. We also use the number of deals between regions as a check in the robustness section but the results remain qualitatively unchanged. We construct a gravity model to explain M&A investments between European NUTS 2 regions and to estimate the average European border effect. Our gravity model is based on annual panel data and is similar to that of Loungani et al. (2002), who use gravity panel data on FDI. In a similar vein, Portes and Rey (2005) explain cross-border equity investments, and di Giovanni (2005) cross-border M&A transactions. Following the approach of Silva and Tenreyro (2006), we estimate a Poisson regression and a pseudo maximum likelihood estimator (PPML). We introduce a panel estimation with random effects to capture the border effect's development over time:

$$\ln\left(y_{t,i,j}\right) = \ln\left(GDP_{t-1,i}\right) + \ln\left(GDP_{t-1,j}\right) + \ln\left(Dist_{i,j}\right) + c_{i,j} + \mathbf{X}_{t,i,j} + \varepsilon_{i,j} \tag{1a}$$

$$\ln\left(y_{t,i,j}\right) = \ln\left(GDP_{t-1,i}\right) + \ln\left(GDP_{t-1,j}\right) + \ln\left(Dist_{i,j}\right) + \mathbf{C}_{i,j} + \mathbf{X}_{t,i,j} + \varepsilon_{i,j} \tag{1b}$$

In both equations, the dependent variable $y_{t,i,j}$ represents the aggregated amount of investments made in period t by acquirers located in region i into region j. As stated before, all investment flows are measured by the sum of transaction value. We use inflation-adjusted values for all variables reported in currency. The variable *GPD* contains the gross domestic product on NUTS 2 level for acquirer region i and target region j in the year before to avoid endogeneity. The macroeconomic data is taken from the EUROSTAT regional statistics website. EUROSTAT provides GDP data on NUTS levels only from 1995. We fill in our own estimates for the years with missing data. Therefore, we extrapolate regional GDP data based on the individual growth rate of each NUTS region in the years 1995 to 2007 to the missing years, 1990-1994. Restricting our sample to the years 1995-2007 does not change our results. In the robustness checks section we also include GDPs of neighboring regions to control for spillover effects. $Dist_{i,j}$ measures the distance in kilometers between the centroids of each region i and j. Similar to McCallum (1995) and Anderson and Wincoop (2003) we use indicator variables to control for border effects. In equation (1a) we use a single indicator variable $c_{i,j}$ to control for cross-border transactions in the pooled sample. Equation (1b) allows the cross-border coefficients to vary over time; we use $C_{t,i,j}$ as a matrix of annual indicator variables.

$X_{t,i,j}$ is a matrix of various control variables, mostly on country level. In many gravity model regressions, common borders play an important role, e.g., because of shared cultural or ethnical ties. To control for these effects, we include a dummy variable that takes the value of one if two countries share a common border (*SameBorder*). Other cultural differences can have significant impact on investment decisions. We control for cultural effects in our sample using a cultural distance measure (*CultDist*) constructed by Kogut and Singh (1988) which in turn is based on Hofstede's dimensions of national culture (see Hofstede 1980, 1983). Different legal systems might influence the decision to buy another firm; following La Porta et al. (1998 and 2007), we include a dummy variable based on similar legal systems (*SameLaw*) between countries. As the overall business cycle might influence M&A investment behavior, we also include a variable that contains the one-year stock index performance *StockIndexPerf* = I_t / I_{t-1}. To account for possible other national differences we include country fixed effects in most of our analyses.

III. Empirical Results

Regression models

As a first benchmark and to allow for comparison with some of the older literature we start with an ordinary least squares (OLS) panel regression to estimate equation (1a) (for OLS gravity models see, e.g., Stein and Daude 2007). Table 2 contains the regression results of our gravity model with different specifications. We find highly significant positive effects of both, target and acquirer regions' GDPs (*log(GDP_Tg) and log(GDP_Aq)*, respectively) as well as a significant negative effect of distance (*log(Dist)*) on investment flows between European regions (Table 2, model 1). Typically for gravity models, these coefficients are highly significant; this fact can also be observed in analyses of exports and foreign direct investments. Model 1 includes the main dummy variable that is one if the two regions are located in two different countries and zero otherwise (*CrossBorder*). Its coefficient shows a significantly negative impact of national borders on total investments between regions. Even though the model controls for distance between and economic activity within each region, two regions of different nationality show significantly lower investment flows than region pairs within one country.

Table 2: Gravity Models - Results

This table reports the regression results of various gravity models according to equation (1a). The dependent variable in all models is the log of total investment value *log(TgValue)* directed towards the target region. Model 1 is an ordinary least squares (OLS) panel regression. Models 2 to 7 are Poisson models estimated by pseudo maximum likelihood method (PML). Exogenous variables start with the gravity components, i.e., GDP of target and acquirer region in the prior year and distance measured by the centroid distance between target and acquirer region. Additional variables are: *CrossBorder* as an indicator variable for region pairs that cross national borders, AR(1) are autoregressive components that contain the log of investments in the previous period based on the region pair (*individual*) or all acquiring regions (*group*) into the respective target region. We additionally control for similarity of legal systems (*SameLaw*), the cultural distance between nations (*CultDist*), for investments in adjacent countries (*SameBorder*), and for one-year stock index performance (*StockIndexPerf*). Model 7 uses our US-Sample for the same estimations, except for the *CultDist* and *SameLaw* control variables. All panel regressions are random effects models.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Estimation Method:	*GLS*	*PPML*	*PPML*	*PPML*	*PPML*	*PPML*	*PPML*
log(TgValue)=	*EU*	*EU*	*EU*	*EU*	*EU*	*EU*	*US*
log(GDP_Tg)	0.019**	0.732**	0.744**	0.627**	0.916**	0.912**	1.183**
	(22.72)	(29.10)	(29.42)	(24.08)	(30.49)	(30.45)	(36.96)
log(GDP_Aq)	0.023**	0.914**	0.927**	0.901**	1.315**	1.320**	1.621**
	(27.89)	(36.33)	(36.65)	(36.09)	(43.55)	(43.92)	(51.22)
log(Dist)	-0.032**	-0.732**	-0.737**	-0.715**	-0.589**	-0.519**	-0.457**
	(-26.84)	(-18.66)	(-18.71)	(-18.63)	(-16.18)	(-14.50)	(-20.83)
CrossBorder	-0.071**	-1.672**	-1.685**	-1.629**	-1.558**	-1.494**	-0.679**
	(-21.54)	(-21.30)	(-21.38)	(-21.19)	(-21.00)	(-15.99)	(-9.97)
AR(1) individual			-0.036**				
			(-7.30)				
AR(1) group				0.056**	0.023**	0.024**	0.013**
				(13.72)	(5.13)	(5.27)	(4.09)
SameLaw						0.161	
						(1.90)	
CultDist						-0.178**	
						(-4.50)	
SameBorder						0.188*	0.090*
						(2.45)	(2.20)
StockIndexPerf						1.071**	0.598**
						(15.26)	(3.57)
const	-0.125**	-15.864**	-16.076**	-14.946**			
	(-8.57)	(-41.54)	(-41.91)	(-39.01)			
year FE	No	No	No	No	Yes	Yes	Yes
TgNation FE	No	No	No	No	Yes	Yes	Yes
AqNation FE	No	No	No	No	Yes	Yes	Yes
N	639812	639812	639812	639812	639812	633233	508793
groups	37636	37636	37636	37636	37636	37249	29929
Chi squared	4827.42	5120.21	5160.19	5410.72	7672.05	7739.98	18219.72

All panel regression contain random effects, ** and * denote statistical significance at 1% and 5% levels, respectively

OLS estimations could be biased because of the dependent variable's heavily skewed distribution. Therefore, we estimate the same equation with a Poisson regression in model 2. In general, coefficients become larger in our Poisson estimation, e.g., the *CrossBorder* coefficient changes from -0.071 to -1.682. Notwithstanding, both OLS and Poisson models result in the same signs and similar significance levels for all coefficients. National borders significantly impede M&A transactions from one region to another.

Some motivations for M&A investments in specific regions may not be captured by local GDP and national control variables. There could be regional features such as a region's specific industrial composition or public infrastructure like ports or airports. Finally, investments might be triggered not only by economic factors but also by behavioral motives like herding. Competitors that buy firms in a certain region could trigger an acquirer's decision to invest in that region, too. Both arguments suggest the use of a control variable that takes into account other firms' investment in a specific region. We do so by inserting two autoregressive components in our model. In model 3, we add a first order autoregressive term (*AR(1) individual*) which contains the individual region-pair investments of the previous period. To control for all investments in a specific region, we include an autoregressive term in model 4 that contains investments from all acquirer regions into the respective target region in the previous year (*AR(1) group*). The individual region-pair coefficient in model 3 is significantly negative, which does not confirm the notion that herding plays a large role in investment decisions. When taking all investment source regions into account in model 4, the AR-coefficient gets significantly positive (0.056), supporting the argument that there might be other regional features that trigger investments in a specific region. Including these controls leaves the cross-border variable almost unchanged (from -1.672 to -1.685 for individual and to -1.629 for group autoregression).

Of course many other factors could influence these results from the standard panel regression. In model 5 we introduce year-fixed effects to control for the business cycle. Also, country-specific price, supply and market characteristics could distort our estimations. Anderson and van Wincoop (2003) derive a unique control variable from price indices to correct for multilateral resistance. As shown by Feenstra (2002) and Anderson and Wincoop (2003), country-specific fixed effects also give consistent estimates. We use country fixed effects for home and host country separately as in

Redding and Venables (2000), Rose and Wincoop (2001) and others. Model 5 incorporates both year and country fixed effects. Our results are robust against these changes. Consistent with the findings of Anderson and Wincoop (2003), the cross-border coefficient gets smaller, from -1.629 to -1.558, and remains highly significant. There are some changes in the outcomes for local GDP and distance, but the coefficients remain qualitatively unchanged and stay highly significant.

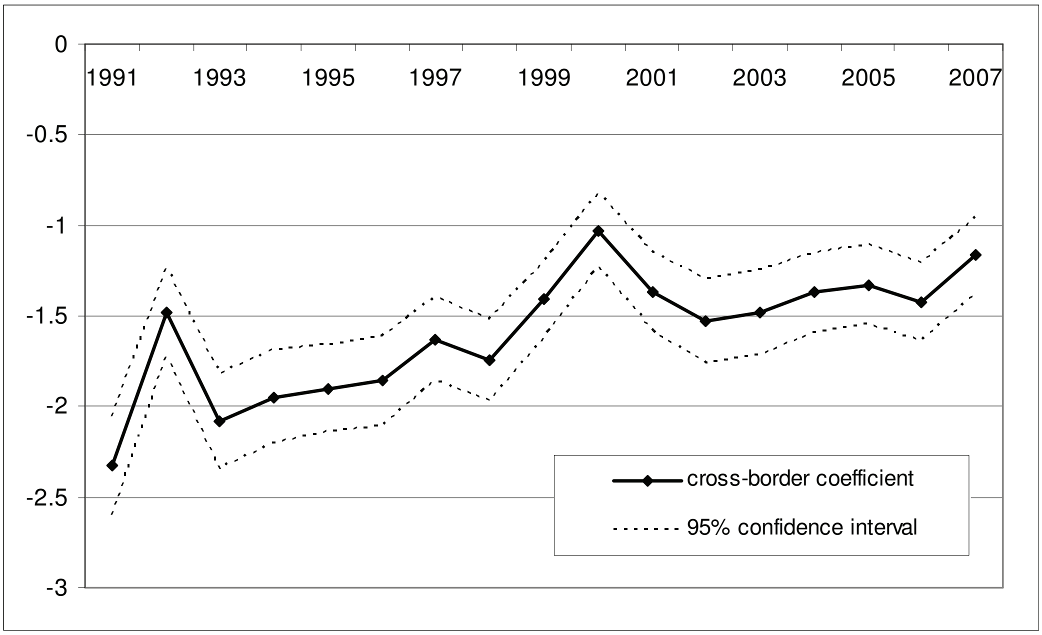
In model 6, we include control variables for legal, cultural and geographical aspects that might influence the willingness of firms to buy others in different regions. The indicator variable *SameLaw* equals one when the two regions are in different countries with similar legal systems, as specified in La Porta et al. (1998, 2007). As expected, the coefficient is positive, albeit only significant at the 10 percent level: having similar law systems encourages transactions between regions in different countries. The coefficient for *CultDist*, the variable that measures the difference in the national cultural index between two countries according to Kogut and Singh (1988), shows a highly significant and negative coefficient. The larger the cultural distance between countries the lower the level of M&A transactions between two regions—even in addition of physical proximity. Also of interest is the *SameBorder* variable which equals one if the countries of the respective regions share a common border. The coefficient is positive and significant on a 5 percent level: even after controlling for distance and other control variables, a common border enhances cross-border M&A transactions. This is a common finding in the trade and FDI literature. To control for share prices as a driver of M&A activity as suggested by Shleifer and Vishny (2003) we control for the one-year stock market performance (*StockIndexPerf*) of the Dow Jones STOXX Europe Index for Europe. The cross-border coefficient, however, does not change qualitatively with the introduction of these control variables. Interestingly, the introduction of these control variables reduces the *CrossBorder* coefficient only to a small extent from -1.558 to -1.494, while the significance level is somewhat reduced but remains very high (z-value of -15.99). Even after controlling for legal and cultural differences, for common borders, and even after including time and country fixed effects to capture other idiosyncratic country facts, transactions between regions in different countries remain significantly impeded. In all specifications we find significantly negative coefficients for the cross-border indicator variable on a level of roughly -1.6 throughout all Poisson models. Borders have a constraining effect on M&A investment flows between European regions.

Development over time

In addition to this static result we are interested in the development of the border effect over time. One of the major aims of the European Union is to form an increasingly integrated market; we expect to see a lessening of the border effect over time. In order to test this hypothesis, we modify our regression model and allow for annual changes in the cross-border coefficients according to equation (1b). Figure 3 shows the annual coefficients for the cross-border indicator (black line) together with the coefficient's 95 percent confidence interval (dotted line). The underlying regression equation is the same as in Table 2, model 6. Except for the now time-varying cross-border components all other coefficients show similar size and significance (Table not reported). The graph starts in 1991 since we include a first-order autoregressive term in the regression equation as explained above.

Figure 3: Cross-Border Impact over Time

This figure shows the annual coefficients of the cross-border indicator variable (black line) together with their 95 percent confidence interval (dotted line) based on equation (1b) with all exogenous variables from Table 2, model 6. The graph starts in 1991 as we include a first-order autoregressive term in the equation for dynamic control.



Overall, Figure 3 depicts an upward trend during the observation period from roughly -2.0 in the early nineties to about -1.25 in 2007, in line with the hypothesis of an increasingly integrated European market. On average the level corresponds to the single component estimate of -1.494 in Table 2, model 6. Integration seems to proceed at a rapid pace until the year 2000, where there is an absolute peak with a border coefficient of almost -1.0. Since the Euro as a common European currency was introduced in 1999 in 11 of the 15 countries in our sample, this seems to be a logical development. The strong increase in absolute terms of the border coefficient after 2000—in 2002 and 2003 the coefficient drops again to levels of around -1.5—comes as a surprise. Borders in Europe do matter more again. Manchin (2004) analyses EU-transactions until 2001 and does not find a Euro-effect, too. The coefficient remains at around -1.4 until 2006; only in 2007 it does change to -1.17, the second highest observation after 2000. If it was not for the year 1992, the confidence intervals in the early nineties would not overlap with those at the end of our observation period. Statistically, this graphical interpretation is similar to a mean-comparison t-test and suggests a significant decrease in border effects over time. However, when comparing an eleven year difference between 1996 and 2006, the confidence intervals do overlap. In other words, the coefficients in 1996 and 2006 do not differ in a statistically significant way—the border effect does not change between these years. That draws a rather disappointing picture of European integration, at least with regard to the market for corporate control.

IV. EU and US in Comparison

Also drivers not related to European borders and other control variables might hinder cross-border M&A transactions in Europe. For instance, even in a very homogenous market, acquirers might have preferences for merging with firms in a given area: the cross-border coefficient might be different from zero even in a world without direct hindrances. In other words, the observed level of the cross-border variable has to be interpreted. Unfortunately, we are unable to control for these other kinds of firms' local preferences. In an attempt to approximate a basic level of 'border effect' in a largely homogeneous market, we turn to the US.

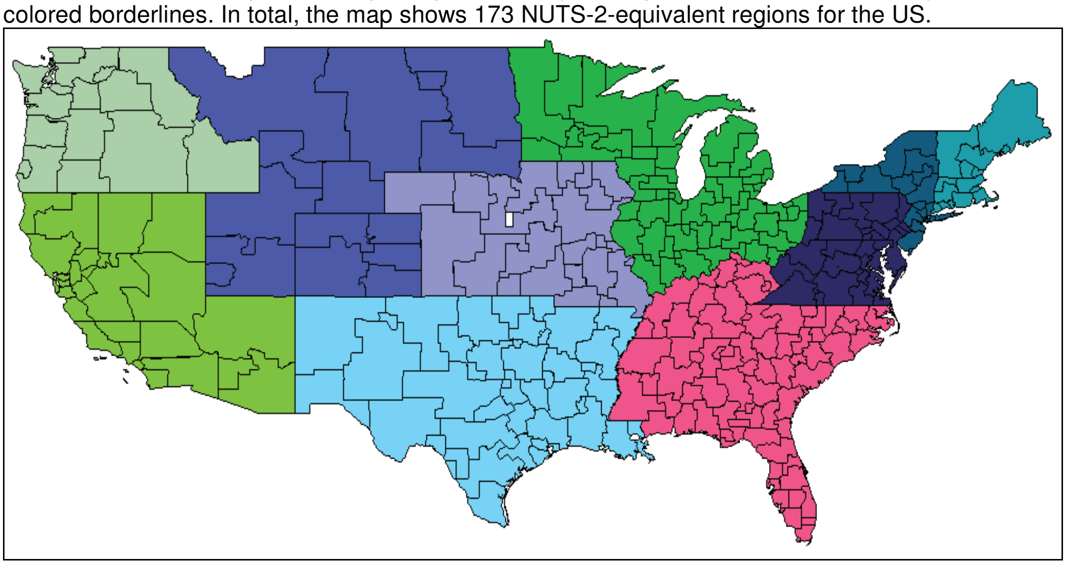
The US and the European Union are similar in geographic and economic size. We make use of one of the most important differences, namely the fact that the US is a single nation with a single language and a largely harmonized regulatory framework.

Some differences in regulations concerning changes in corporate control do exist between states (Bebchuk and Cohen 2003). We nevertheless regard the US as a fairly homogenous market and use it as a benchmark for the EU with its distinct member states. Even when 'weak' borders in the US, e.g., due to geographical industry composition or local preferences might exist, estimates of 'border coefficients' within the US should reflect the maximum level of integration in Europe. The difference between the actual cross-border coefficient in the EU and the quasi 'cross-border' coefficient in the US signals the potential for further EU harmonization.

In order to maximize comparability, we use data on M&A transactions that take place within the contiguous 48 US states (excluding Alaska, Hawaii and others) and apply the same filter criteria as for the European sample. To use geographical entities that are roughly equivalent to European states we define 'quasi-borders' within the United States along the borders of the ten Federal Regions defined by the Office of Management and Budget (OMB 1974). Figure 4 shows the OMB-regions as differently colored areas. Of course there is some arbitrariness in this approach. Nevertheless, OMB borderlines have many characteristics similar to European borders, e.g., borderlines along natural obstacles (rivers Mississippi in the US and Rhein in the EU; the Rocky Mountains and the Alps, etc.).

Figure 4: OMB Regions and NUTS 2 Equivalents

This map of the United States shows the ten standard Federal Regions that were established by the Office of Management and Budget (OMB) Circular A-105, 'Standard Federal Regions', in April, 1974. The OMB-regions are highlighted by the differently colored areas. Additionally, the map contains all NUTS-2-equivalent regions generated by our algorithm which are drawn by their black colored borderlines. In total, the map shows 173 NUTS-2-equivalent regions for the US.



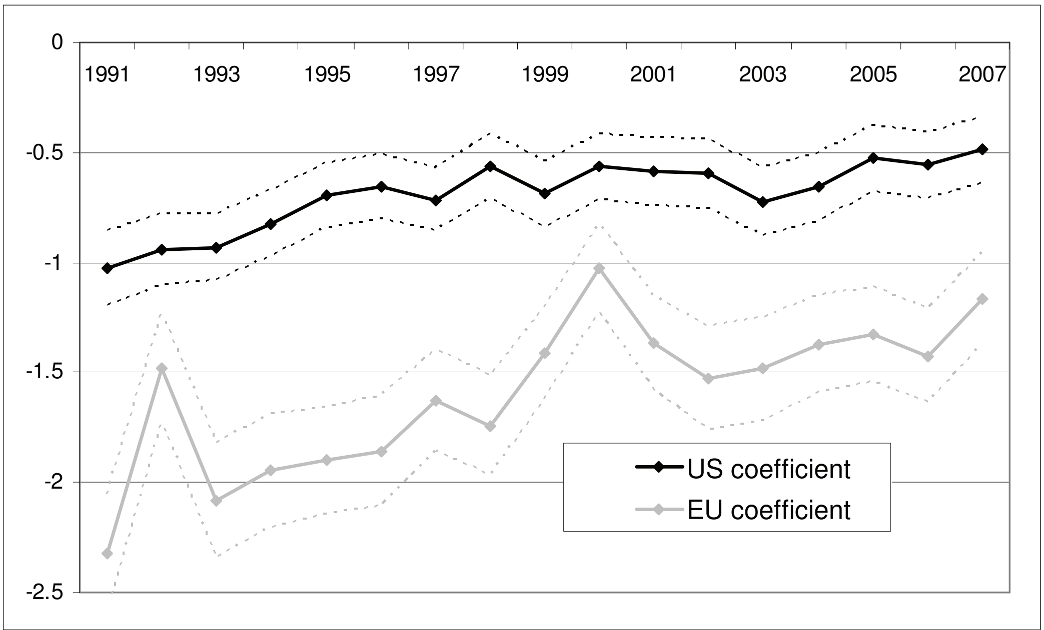
Since US statistics are not reported according to the European NUTS classification standard which we would need to make results comparable, we use county data, which is considerably more detailed than NUTS 2 (more similar to NUTS 3 level) and construct artificial NUTS 2 regions within the US. We apply an algorithm that first sorts all county data by population in the year 2000 in ascending order. We then take the first county and combine it with an adjacent one that has a population below maximum classification requirements (as in NUTS 2 definition) and a location within the same OMB-region. If more than one county is eligible, we take the one with shortest centroid-to-centroid distance. Finally, we merge the two counties and put them back into the pool. We repeat this algorithm until no further county can be matched to any other. In the end, we receive a total of 173 NUTS 2 equivalent US regions in 10 OMB-regions, as compared to our European base-case with a total of 194 NUTS 2 regions in 15 nations. Figure 4 shows a map of all NUTS 2 equivalents; we drop seven counties due to missing data, which explains the small white spots in the map in Nebraska and Montana. We augment the US sample with data on local GDP for our constructed NUTS 2-equivalent regions (inflation-adjusted with base year 2000), distance, and all other relevant control variables. For stock index performance, we use the Dow Jones Industrial Average Index.

We then run a gravity regression similar to model 6 in Table 2 except that we exclude the control variables for legal similarities (*SameLaw*) and the cultural distance (*CultDist*). The coefficients for the control variables show qualitatively the same results as in the European sample. However, the border effect turns out to be significantly smaller. In the single-regressor model according to equation (1a), the significant cross-border coefficient amounts to -0.679 (model 7 in Table 2) which is less than half the size of the European cross-border coefficient of -1.494. Figure 5 displays the cross-border variables' time-varying coefficients (solid black line for the US, solid grey line for the EU). We also show their respective 95 percent confidence intervals (dotted lines). Both graphs are taken from estimations with the same covariates as in Table 2, but with time-variant cross-border dummies, according to equation (1b).

The most striking difference between Europe and the US is the level of the cross-border coefficients. Throughout our observation period, the difference between the US and the EU remains roughly the same over time. The distinctive peak in cross-border activity in Europe in 2000 is not visible in the US analysis. Interestingly, our previous finding of an increasing (decreasing) integration (coefficient) over time exists in the US sample as well: the US in the early 1990s also shows lower levels of cross-border transactions with subsequently higher levels towards the end of our estimation period. This leads us to the conclusion that the lessening impact of national borders in Europe is more likely a general economic phenomenon than the result of deliberate integration measures. Even more striking; with a first-three-years-average of -1.963 and a last-three-years-average of -1.306 the European border effect decreases by only 33.5 percent while the same ratio for the US reveals a decrease of 45.8 percent.

Figure 5: Comparison of EU and US Border Effects

This graph displays the cross-border variables' time-varying coefficients (solid black line for the US, solid grey line for the EU). We also plot their respective 95 percent confidence intervals (dotted lines). Both graphs are taken from estimations with the same set of exogenous variables as in Table 2, but with time-variant cross-border indicators. In the US-estimation we exclude variables for legal similarities (*SameLaw*) and cultural distance (*CultDist*).



One reason for the enduring difference between the US and the EU could be differences in the respective geographical industry compositions. This is not likely to be the case: In both EU and US sample there are about 60 percent intra-industrial and about 40 percent diversifying M&A transactions, when measured on the two-digit SIC level. Given the homogeneous market in the US, industries in the US tend to be more clustered than the EU with its history of separate states (Midelfart-Knarvik et al. 2000). Transactions, especially intra-industry ones, have a higher likelihood to take place locally in the US since there are more choices available locally. Therefore, all other things equal, one should expect more local activity in the US and less cross-border deals than in the EU, which in turn would argue in favor of a stronger 'border' effect in the US.

How much is Europe lagging behind the US?

We are interested in the quantitative meaning of the coefficient's difference between the US and Europe: how much would the transaction pattern within the EU have to change to resemble US regional border effects? Coefficients in Poisson gravity equations do not lend themselves easily to interpretation; we use different measures to gauge a likely range of necessary changes in Europe. First, we compare cross-border investments as a fraction of all transactions measured by the total investment value between Europe and the US. In Europe only about $\varphi_{EU} = 30\%$ of all transactions are cross-border while in the US about $\varphi_{US} = 68\%$ cross an OMB-region border. Accordingly, with a fixed level of domestic M&A transactions Europe would need about $\frac{\varphi_{US}(1-\varphi_{EU})}{\varphi_{EU}(1-\varphi_{US})} = 4.96$ times the current level of cross-border activity to match the US level.

A more elaborate approach is the use of gravity model estimates. Indicator variables are interpreted conditional on the mean of the underlying distribution which in this case can be misleading since our sample's distribution is heavily skewed. Therefore, we interpret our findings cautiously. A standard interpretation of our estimates uses the cross-border-coefficients' antilog with $e^{(-1.494)} - 1 = -77.6\%$ for Europe and $e^{(-0.679)} - 1 = -49.3\%$ for the US (see Halvorsen and Palmquist 1980). After controlling for distance, GDP, etc. our model suggests that European cross-border region pairs on average have 77.6 percent less transaction volume than domestic ones ($\varphi_{EU} = 22.4\%$), while in the US the volume is 49.3 percent less ($\varphi_{US} = 50.7\%$). These figures suggest that Europe needs $\frac{\varphi_{US}(1-\varphi_{EU})}{\varphi_{EU}(1-\varphi_{US})} = 3.6$ times more cross-border deals to reach the US level.

As a third way of retrieving an answer, we counterfactually adjust our European data to construct a hypothetical sample that, by design, results in similar estimates as in the US. We leave the domestic European transaction volume constant and increase all European cross-border investment flows by multiplying their original values. We do that until the confidence intervals of the EU coefficients include the US estimate of the year 2007. As set-up, we use model 6 in Table 2 for the EU15 sample and the same equation in the US (except for the dummies *SameLaw* and *CultDist*) as benchmark. As a result, European cross-border M&A transactions would have to increase by 7.6 times (not reported). For instance, in 2007 an increase of cross-border activity from 36.2 percent to 81.1 percent of total transaction volume would be needed in Europe to match US levels.

All mentioned approaches are counterfactual and do imply, among other things, that M&A opportunities in foreign regions were available, firms had sufficient access to fund these activities, and that domestic activity stayed constant. It is probably more likely that cross-border transactions might substitute domestic ones. In another approach, if we hold the total amount of transactions as constant, 43 percent of the domestic volume in the EU sample would have to become international transactions to match the US figures leading to a 92 percent increase in international transactions.

There is a lot of uncertainty in estimating the increase needed for European cross-border transactions to measure up to US levels. However, all three methods result in multiplier estimates of at least 3.6 times as much cross-border activity in Europe, given the actual domestic transaction volumes. Even when holding constant the total amount of transactions, more than 43 percent of current domestic transactions would have to become international ones to match US data.

IV. Robustness Checks

To ensure the validity of our results, we perform a variety of robustness checks which are reported in the following section. Our main findings about the development of the cross-border coefficient hold through all settings.

Econometric specification

We use Poisson regression models to meet the characteristics of our endogenous variable and to be comparable to other recent papers that have used a similar approach. One shortcoming of Poisson models, however, is the assumption that the conditional mean should be equal to the conditional variance. To account for possible over-dispersion, we run our model additionally with a negative binomial regression model, which allows the variance to differ from the mean (e.g., see Hausman, Hall and Griliches (1984)). The estimates of the negative binomial regression for the European sample are reported in model 1 of Table 3. While basically all control variables seem to be robust with regard to the Poisson model, we find that the cross-border coefficient in the negative Binomial model tends to be slightly higher in absolute terms in most regressions. We find that the Poisson regressions estimate the impact of borders in Europe on M&A activity somewhat conservatively.

Table 3: Robustness Check Regressions

This table reports various robustness regression based on different variations in model set-up and design. Model 1 shows the results of a negative binomial regression as opposed to our Poisson regression. Model 2 and 3 report Poisson regressions with a dependent variable consisting of the number of transactions instead of the value of transactions. The difference between model 2 and 3 is the underlying sample which, in the first case, consists of the sample used in the analyses before, while model 3 additionally contains transactions with unknown deal price. In model 4 we integrate a spatial lag variable to control for investments surrounding the target region. In model 5 we include a spatial filter in the regression equation as suggested by Griffith (2003) to control for spatial autocorrelation.

	(1)	(2)	(3)	(4)	(5)
Estimation Method:	*NegBin.*	*PPML*	*PPML*	*PPML*	*PPML*
Dependent:	*Value*	*Obs_Value*	*Obs_Total*	*Value*	*Value*
log(GDP_Tg)	1.010**	0.985**	0.972**	0.905**	0.522**
	(34.54)	(28.64)	(44.66)	(30.14)	(12.65)
log(GDP_Aq)	1.626**	1.563**	1.475**	1.327**	1.385**
	(66.96)	(55.44)	(78.65)	(44.06)	(44.58)
log(Dist)	-0.322**	-0.302**	-0.370**	-0.505**	-0.528**
	(-27.60)	(-21.75)	(-35.16)	(-14.25)	(-14.42)
CrossBorder	-1.792**	-1.743**	-2.193**	-1.531**	-1.569**
	(-24.46)	(-20.32)	(-40.36)	(-16.34)	(-16.09)
AR(1) group	0.077**	0.075**	0.042**	0.022**	0.012**
	(10.88)	(8.99)	(8.50)	(4.98)	(2.59)
SameLaw	0.125	0.114	0.330**	0.166*	0.136
	(1.64)	(1.27)	(6.53)	(1.97)	(1.55)
CultDist	-0.189**	-0.185**	-0.098**	-0.177**	-0.196**
	(-5.43)	(-4.53)	(-3.79)	(-4.47)	(-4.82)
SameBorder	0.418**	0.410**	0.461**	0.205**	0.184*
	(5.92)	(4.96)	(10.22)	(2.67)	(2.30)
StockIndexPerf	1.188**	1.115**	0.583**	0.858**	0.966**
	(10.25)	(8.26)	(6.72)	(3.17)	(13.68)
Spatial lag on dep.				0.062**	
				(5.59)	
year FE	Yes	Yes	Yes	Yes	Yes
TgNation FE	Yes	Yes	Yes	Yes	Yes
AqNation FE	Yes	Yes	Yes	Yes	Yes
Spatial Filter	No	No	No	No	Yes
N	633233	633233	633233	633233	633233
groups	37249	37249	37249	37249	37249
Chi squared	15942.11	11236.88	19045.4	7739.38	7730.92

All panel regression contain random effects, the asteriks ** and * denote statistical significance at 1% and 5% levels, respectively

Large versus small transactions

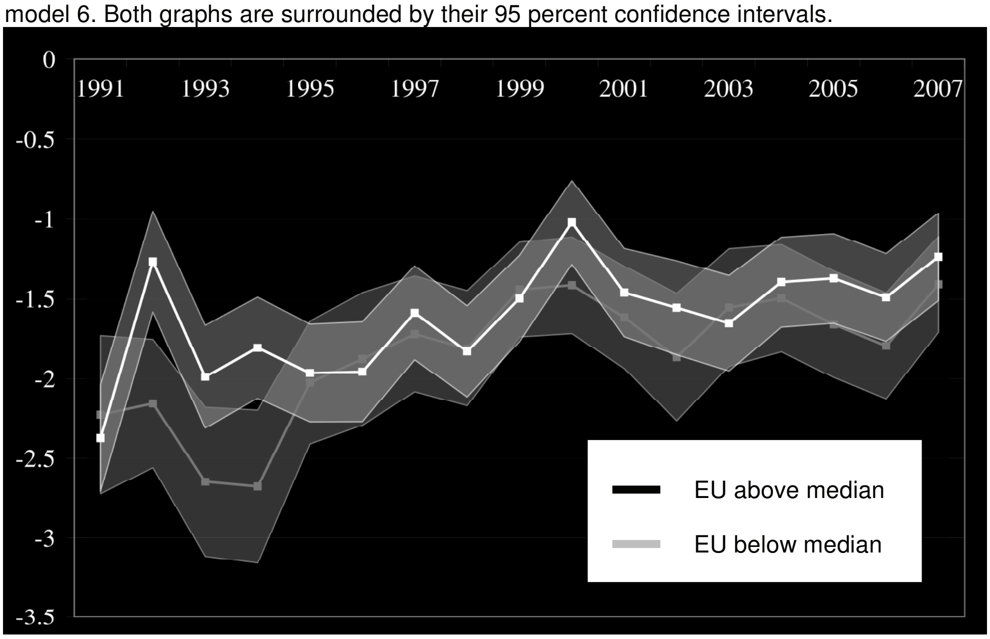
In the analyses so far we use the volume of transactions between two regions as the dependent variable. This might bias our results towards large transactions. As a robustness check we run all regressions with the number of transactions between two regions as the explanatory variable, i.e., weighing all transactions equally. In Table 3, model 2 and 3 report a Poisson regression with the dependent variable consisting of the total number of reported transactions. The difference between model 2 and 3 is the underlying sample, which in the first case contains all transactions of our previous sample (*Obs_Value*), whereas model 3 additionally contains all transactions with unknown deal price (*Obs_Total*). Due to our initial sample selection these transactions were excluded. Compared to our previous coefficients, the cross-border coefficient in model 2 has a higher absolute value which suggests a stronger impact of borders. With a coefficient of -2.193, model 3 reflects an even stronger border effect due to inclusion of transactions with unknown deal value. Most of these deals are transactions by privately owned companies, which tend to be smaller than publicly listed companies and undertake smaller transactions. As expected, borders matter less for larger companies than for smaller companies.

Furthermore, we are interested whether large transactions are affected differently by European borders than are the small transactions in our original sample. To test for different influences in the small and medium sized enterprise segment, we split our data into two subsamples of below and above median target size as measured by the target's market value. For the European sample the median size is about 12.0 million Euro (the median US transaction is about 28 million USD). Since there is no significant median variation over time, we use the full sample median as a fixed threshold. We run the same regressions as before (as in Table 2, model 6); Figure 6 shows the sample of small transactions vs. large transactions.

The graphs display time-varying coefficients of the cross-border indicator together with their respective 95 percent confidence intervals. Both, the coefficients of the below-median sample and the coefficients in the above median sample display an increase over time, i.e., a lessening impact of borders for M&A transactions in the EU. In the majority of years, the graph of smaller transactions is below the graph for the larger transactions, i.e., for the larger transactions borders matter less. Perhaps surprisingly, the two graphs are not significantly different from each other for most of the years (the same holds true for the US, figures not reported).

Figure 6: Firm Size and Cross-Border Effects in the EU

We split our sample into two subsamples along the median target market value. This graph shows the cross-border coefficients for the two subsamples of below (grey line) and above (black line) median target size. The estimations contain all control variables of Table 2, model 6. Both graphs are surrounded by their 95 percent confidence intervals.



Spatial interdependence

M&A activity in one region could be influenced by the characteristics of neighboring regions or the M&A transactions taking place there. This might bias our findings since small regions that are not of interest in itself might attract investments because of their proximity to other, more attractive regions. The problem is somewhat mitigated in comparison to Greenfield FDI, because for an M&A investment a suitable firm has to be in place. We nevertheless account for possible spill-over effects by two measures, first, by introducing ‘spatial lags’, and second, by controlling for spatial autocorrelation (SAC). None of this changes our main findings. Spatial lag variables are control variables that account for the amount of investment in neighboring regions to explain the amount of investment within one region, similar to time lags in time series analysis. As within the SAC approach, there are several ways to construct a spatial weight matrix (see Getis and Aldstadt 2004); one frequently used form is the contiguity matrix that weighs adjacent regions with one and all other elements with zero. A more sophisticated form uses inverse distance weights for all other regions. Following the literature, we include investments in all regions weighted with their inverse squared distance into the spatial lag variable for each region. Model 4 in Table 3 shows the results. We include the same control variables as before. The spatial lag variable is positive and significant, thus showing indeed transactions in neighboring regions influencing activity in a given region positively. Other controls remain largely unchanged. The cross-border coefficient is -1.531 and highly significant, remaining slightly bigger than before with a coefficient of -1.494. Thus, our main findings do not change: borders within the EU remain a barrier for M&A transactions even when we account for the attractiveness of neighboring regions.

Estimations using spatial lag variables could be biased because some dependency is introduced into the observations (see Ord 1975). We therefore turn to another robustness check to control for spatial spillovers: the spatial autocorrelation technique. One recent and commonly used SAC approach is an eigenvector-based spatial filtering technique promoted by Griffith (2000, 2003). This procedure creates a so-called spatial filter by simply adding a set of control variables to the regression equation. This approach does not need assumptions about the dependent variable's distribution and thus can be applied easily to Poisson regression models. The method builds on an eigenvector decomposition of a spatial weight matrix which contains distance information between all regions. As in the spatial lags case above, we construct a spatial weight matrix **W** using inversed squared distances $w_{i,j} = \left(1/d_{i,j}^{2}\right)$ between regions *i* and *j* for all $i \neq j$. This common form allows for a quadratic decline in distance. Following Griffith (2003) we transform **W** into $\overline{\mathbf{W}} = \left(\mathbf{I} - \mathbf{1}\mathbf{1}^{T}/n\right)\mathbf{W}\left(\mathbf{I} - \mathbf{1}\mathbf{1}^{T}/n\right)$ where **I** is the *n x n* identity matrix and **1** is the *n x 1* unit vector before conducting the eigenvector decomposition. This allows for the decomposed vectors to be ordered by their spatial correlation, which is measured by Moran's I coefficient (see Griffith 2003). Through decomposition of $\overline{\mathbf{W}}$, we derive as many orthogonal eigenvectors as there are NUTS regions in our sample, of which we select all eigenvectors with an absolute Moran's I of at least 0.2 and add them as control variables to our regression model. Altogether 89 (orthogonal) eigenvectors match this criterion. Model 5 in Table 3 reports the regression results of a Poisson regression with inclusion of this spatial filter. The coefficient of our cross-border indicator becomes a bit bigger in absolute terms (-1.569 as opposed to -1.494 in Table 2, model 6). M&A investment in Europe and the border coefficient are to a small part driven by the influence of other surrounding regions. Similar to the findings of Bloningen et al. (2007), who analyze FDI from the US, our results do not change qualitatively when we control for spatial autocorrelation in our model.

V. Conclusion

Our study focuses on European integration in the market for corporate control. We analyze M&A transactions that took place between 1990 and 2007 in the 15 countries that formed the EU in 1995. Basic units of analysis are the European NUTS-2 regions with up to 3 million inhabitants. We construct a gravity equation model to explain transaction volumes between regions, while controlling for a variety of possible influences known from the literature such as distance and local economic activity in acquirer and target regions, cultural and legal differences, whether countries have a common border, as well as time and country fixed effects to control for omitted variables. Our interest focuses on a cross-border indicator variable that equals one when a transaction crosses a national border and zero otherwise. We find a strongly negative, highly significant and robust border effect under all specifications and estimation techniques in our pooled regressions. In order to track the impact of borders on European M&A transactions over time we introduce a dynamic Poisson panel analysis and find a decreasing restraining effect induced by borders in our observation period. There is, however, a notable increase in the importance of borders after the year 2000, when the Euro was introduced and at the same time the dot-com bubble burst. Between 1997 and 2007 we find no significant change in the border effect.

Measuring the impact of European integration on the border effect in M&A transactions is only a part of the picture. After all, the effect of declining relevance of borders on foreign direct investments (FDI) is ambiguous. On the one hand, when FDI are done by firms that exploit local comparative or absolute advantages, such as cheaper labor costs or better market access, lower barriers between countries would lead to more international investments within the region (Baldwin et al. 2003; Puga and Venables 1998; Frey and Hussinger forthcoming). On the other hand, one could argue that due to integration cross-border FDI might loose necessity since markets are more easily accessible. Bjorvatn (2004) shows, however, that economic integration encourages cross-border investments. Furthermore, if FDI are undertaken to avoid tariffs and other transport costs that occur when national borders still matter economically, a reduction of FDI should occur along with increasing regional integration (Altomonte 2007; Markusen 2002; Markusen and Venables 1998). The European Union (earlier: European Community) had already abandoned intra-European tariffs by 1968, so, given decreasing borders, one would strictly expect more FDI.

However, there might be other factors beyond our control variables that influence the results. To create a benchmark for European developments, we compare the results with developments within the US, where actual borders between regions never existed. In two steps we divide the US geographically to make it comparable to Europe. Based on counties, we first construct regions that are similar to the NUTS2 regions used in Europe, and second we divide the US in ten different artificial ‘countries’ along the borderlines of the ten Federal Regions as defined by the Office of Management and Budget. Running the same regressions as in Europe, we find M&A transactions in the US far less restrained by borders than transactions in Europe. The overall US ‘cross-border’ coefficient is less than half the European one (-0.7 as opposed to -1.5), but still significantly different from zero. The ‘country’ borders within the US—whose crossing bear almost no additional costs—do restrain M&A investments to some extent. Running the same dynamic panel regressions we find qualitatively the same results as in Europe, namely that the border effect decreases over time. Similarly to European developments, the strongest changes occur in the early 1990s when the US experienced a similar ‘integration’ effect to the one that took place in Europe. In fact, while the European border effect decreases by 33.5 percent during our observation period, the ‘border effect’ within the US decreases by 45.8 percent. Our results show that Europe is integrating at a slower pace than the US in the market for corporate control. This places some doubt on the success of political measures for European integration.

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