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Andre V. Mollick
The University of Texas Rio Grande Valley, andre.mollick@utrgv.edu

Rene Ramos-Duran

Esteban Silva-Ochoa

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Andre Varella Mollick* Rene Ramos-Duran†

Esteban Silva-Ochoa‡

*Dept of Economics and Finance, University of Texas Pan-American, Edinburg, Texas, USA, amollick@utpa.edu

†Industrial Development Secretariat of the Government of the State of Chihuahua, Mexico, renecaesar@yahoo.com

‡Mckinsey & Company, Mexico City, esteban_silva@mckinsey.com

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Abstract

In December 1993, restrictions to foreign ownership across major Mexican economic sectors were abolished. This paper studies output, industrialization intensity, "international infrastructure", and government expenditures on infrastructure as determinants of FDI inflows into Mexican states over 1994-2001. We conduct a "general to specific" estimation strategy across Mexican states. Telephone lines appear to be very important to FDI as their coefficients are around 2.0 in Random Effects Models. Industrialization is also important, with coefficients varying from 0.62 to 0.67. Allowing for endogeneity between FDI and real output, dynamic GMM panels confirm the robust effects of telephone lines on FDI. International infrastructure thus appears more conducive to FDI than domestic infrastructure, such as interstate and secondary roads. With international infrastructure being a major catalyst of FDI inflows into Mexico, we provide support to ongoing conventional wisdom promoting such type of investment.

KEYWORDS: Agglomeration, FDI, Infrastructure, Mexico, Panel Data

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INTRODUCTION

Policy makers are often interested in knowing which forces attract foreign capital into a country or region. Potential candidates include revenue (market size, country's fundamentals) and cost factors (access to foreign markets, regulation, and wages) that affect the profitability of firms. In Mexico, in particular, some states have benefited more than others after two almost simultaneous events: i) the December 1993 foreign investment law that abolished restrictions to foreign ownership in major sectors; and ii) the introduction of NAFTA in 1994. Recognition of the infrastructure role is provided by government sponsored agencies when discussing airport expansion programs and linking these to capital inflows: "Queretaro and Puebla have benefited significantly from international trade and foreign direct investment (FDI) … By promoting modern and innovative airport infrastructure, both states will enhance their ability to attract foreign investment." (Maquila Portal, 2002).

If this reasoning is correct, some questions remain unanswered. Among them are: i) what kind of infrastructure contributes to attract FDI?; and ii) for a given infrastructure level, is the relationship between agglomeration economies (manufacturing activity as share of state GDP) and FDI positive as expected? This article addresses these issues and studies Mexico's recent experience of attracting large sums of FDI: from US$ 5 billion in 1994 to over US$ 10 billion in 1995 and to the unprecedented US$ 27 billion in 2001. Together with China, Mexico has been used as a good attractor of FDI into emerging markets in theoretical work, such as Aizenman (2003). Empirically, the recent survey on Mexico by Kose et al. (2004) documents significant structural breaks in FDI levels in 1993, as well as increases in investment by multinationals in response to NAFTA.

Agglomeration effects do, in fact, complement the microeconomic theory of the firm and transaction costs. The framework by Dunning (1988) generally lacks agglomeration forces and has been called *OLI-theory* due to: O of ownership advantages and firm-specific assets, L of country-specific factors, and I to the internalization of firm's proprietary assets. Applying the OLI-theory to a disaggregated panel of the U.K. food sector, Giulietti et al. (2004) find predominant ownership-specific and industry characteristics with marginal macroeconomic effects. Also using the OLI framework, Zhang (2001) concludes that China's huge market size, liberalized FDI regime, and improving infrastructure are attractive to multinationals.

Several studies have adopted a *regional perspective to FDI inflows*. Based on the uneven regional distribution of FDI in Turkey, Deichman et al. (2003) find that agglomeration, depth of local financial markets, human capital, and coastal access are helpful in attracting FDI. Buch et al. (2003) employ gravity equations to European FDI and Tuan and Ng (2003) explore why Guangdong has been able to continuously attract and mobilize the largest share of FDI flows from Hong Kong into China and especially into its Pearl River Delta region during the 1990s. The process of manufacturing relocation from city core (Hong Kong) to its peripheral locations (the Pearl River Delta) represents a case of spatial development of a *core-periphery system*, as envisaged by Krugman (1991).

A study on Japanese acquisitions in the U.S. by Bloningen (1997) combines the OLI-paradigm with exchange rate effects as elaborated by Froot and Stein (1991). The quality of infrastructure, the degree of industrialization and the level of inward FDI into the market are used by Wheeler and Mody (1992) as agglomeration measures and are found to be major determinants of FDI in U.S. manufacturing firms in the 1980s. Braunerhjelm and Svensson (1996) find disaggregated agglomeration effects, particularly in technologically advanced industries. On cross-country differences, Mody and Srinivasan (1998) contend that U.S. firms are more influenced by domestic market conditions, in contrast to Japanese firms. Head et al. (1995, p. 228) refer skeptically to "the near impossibility of selecting and correctly measuring all of the relevant infrastructure and input price information that affect location decisions. Omitted sources of attractiveness would almost certainly induce a correlation between the error term and the agglomeration variables."

Having this literature as background, several features can be associated with our study. First, we employ a unique data set on Mexican states, thus capturing both time series (1994-2001) and cross-section (22 highest ranking states by FDI amount) effects through panel data models. We do not know of any other study that explores panel data methods on Mexican FDI since Love and Lage Hidalgo (2000) apply the time series approach of Barrell and Pain (1996), while Feenstra and Hanson (1997) and Khawar (2003) study FDI effects into relative wages and productivity in Mexico, respectively. Second, the Mexican territory is divided into six regions that share geographic features, such as the Northern border states and the top 10 state destination of FDI inflows.[1] Third, we formulate the benchmark specification with four independent variables and employ a "general to specific" approach in order to assess the adequacy of the benchmark model. We believe that the residual-based approach mitigates a lot of the uncertainty regarding the "true model specification". We also take into account endogeneity problems between FDI and real output and employ dynamic generalized method of moments (GMM) panels. Fourth, the relationship between agglomeration economies and infrastructure is an important part of this study. In contrast to Cheng and Kwan (2000), Campos and Kinoshita (2003) and Deichman

[1] Several papers put forward a regional approach to FDI but explore reverse causation mechanisms: FDI affecting GDP growth in Sun and Chai (1998) or FDI affecting state exports in Leichenko and Erickson (1997) and Sun (2001). An exception is Sun et al. (2002) who study regional determinants of FDI in China, although they do not handle industrialization intensity.

et al. (2003) who appeal to lagged FDI as agglomeration forces, we introduce explicitly industrialization intensity.

This study attempts to fill the gap of no systematic study on the relationship between FDI and infrastructure for a developing economy. We conduct estimations at Mexican states over 1994-2001, following a “general to specific” strategy. The concentration of telephone lines appear to be very important to FDI as their coefficients range from 1.98 (all states) to 2.02 (all states but DF) in panel Random Effects Models (REM) estimations. Agglomeration is also important, with coefficients varying from 0.62 to 0.67. The article discusses the robustness of these findings along several dimensions, including endogeneity between FDI and real GDP. Under dynamic GMM panels, the effect of telephone lines on FDI increases slightly to around 3.0, contrasting to other explanatory variables whose coefficients do not become statistically significant.

This paper contains five more sections. The next provides an overview of FDI inflows into Mexico and the following one introduces the econometric methodology. Subsequent sections list the variables in this study and present the major results. The last offers conclusions and presents extensions for future work.

THE PATTERN OF FDI INFLOWS INTO MEXICO

FDI inflows into Mexico have experienced a significant upheaval since the commercial openness of the country in the mid 1980s. Original data for this paper are taken from INEGI’s annual *Census of States*, available in hard copies and not electronically. The introduction of a law on foreign ownership restrictions in late 1993 and NAFTA implementation in 1994 suggest the choice of 1994 as the starting date. On the end date of the sample, INEGI’s *Census of States* publications occur with substantial time lags. We thus gathered data from 1994 to 2001, since data for 2002 were not complete at the time of data construction in 2004.

As shown in Figure 1, from 1980 to 1985, Mexico received 1,298 million U.S. dollars (USD) in yearly averages, which incremented almost 200% during 1986 – 1993, following Mexico’s entrance into GATT. The total FDI amount climbed to an average of USD 3,468 million in the 1986 – 1993 period. The FDI boom into the country, however, actually started with the NAFTA signature in 1994, when Mexico received USD 10,657 million of FDI. Almost simultaneously, in December 1993 a Foreign Investment Law abolished restrictions to foreign ownership in major Mexican economic sectors. Under the law, foreign investment is permitted in any sector, unless specifically mentioned as subject to restrictions (Mexico, 2000). Due to the overall economic uncertainty that followed the currency crisis of 1995, the 1994 FDI amount could not be maintained in the following two years. After declining in 1995 and 1996, the trend reverts again and becomes positive during the post NAFTA period climbing to the yearly average of USD 12,912 million.

In addition to this growing trend at the aggregate, the distribution of FDI inflows into Mexico is heavily concentrated across geographical locations. Table 1 offers a general overview of the 1994-2001 period and explains which states are omitted from the analysis due to data construction problems. Just the Distrito Federal (Mexico, DF, the capital) received an average of 60% of the total FDI amount directed to Mexico during the period between 1994 and 2001. If we consider the Estado de Mexico (the political entity that includes the metropolitan zone of Mexico City) as well, this participation rises to 64%.[2] Next to the metropolitan zone of Mexico City, the Border Region of Mexico attracts most FDI across the country, with the six Northern Border states receiving about one fourth of the national total. The remaining states receive just 10%, on average, of total FDI inflows.

As figure 2 shows, the relative importance of Mexico City's Metropolitan Zone declined considerably in the six years period after NAFTA (1994 – 2000). The trend reversed in 2001, in which the unprecedented FDI inflows can be explained by the acquisition of the biggest financial institution of the country (BANAMEX) by U.S. Citicorp.[3]

HYPOTHESIS TESTING AND EMPIRICAL METHODOLOGIES

The basic theoretical mechanism of determining FDI factors across Mexican states includes effects on revenues and on costs of investors, as put forward by Carlton (1983) and Coughlin et al. (1991). The restricted profit function Π_{ij} of firm i in state j can therefore be written as:

$$\Pi_{it} = K_0 X_{1j}^{\alpha 1} \dots X_{mj}^{\alpha m} (\exp \varepsilon_{it})^N \qquad (1),$$

where: X_{sj} = exogenous variables (s = 1,…, m) at state j, K_0, N, α_1, … ,α_m are unknown constants, and ε_{it} is the firm-location specific effect, which is assumed to be independently distributed across i and j and follow the Weibull distribution. The unknown N is the number such that the 1/N th root of the specific effect

[2] The accounting method for FDI inflows into Mexico, DF, could overestimate the amount of funds towards the capital, as flows to Mexico City's metropolitan zone are taken to be flows into the DF. This is a consequence of many national and international firms establishing their headquarters in Mexico, DF.

[3] In May of 2001, Citigroup purchased the Banco Nacional de Mexico (BANAMEX) for USD 12,500 million. This figure represented 47% of total FDI received by the country in 2001 and about 62% of the total received by the Distrito Federal.

follows a Weibull distribution. Taking logarithms of (1) and dividing by N, we obtain:

$$\ln \Pi_{it} / N = K_1 + \sum_k \beta_k \ln X_{kj} + \varepsilon_{it} \qquad (2),$$

where: $K_1 = \ln K_0 / N$ is a constant and $\beta_k \equiv \alpha_k/N$. Firm i locates in state j* provided profits are highest in state j* ($\Pi_{ij*} = \max_j \Pi_{ij}$). The (RHS) of (2) for state j* must exceed that for other states. Panel data regression models will provide elasticities, as our specifications are defined in logarithms aggregated at the state level:

$$FDI_{jt} = \alpha_j + \beta X_{jt} + \varepsilon_{jt} \qquad (3),$$

where j = 1, …, 22 states; FDI_{jt} are the (logarithm) FDI flows into the j-th state, the vector X_{jt} denotes the FDI various potential determinants (in logarithms) listed below, and ε_{jt} are serially uncorrelated errors with zero mean and constant variance. We form six regional groups of j states as follows: all 22 states, all states but DF (21 states), the six U.S.-Mexico border states, top 10 ranked states ("rank") in terms of FDI inflows regardless of geographic region, non-border states (16 states), and non-border states but DF (15 states). We focus on the first four groups, omitting the two groups of non-border states.

Applying the Fixed Effects Model (FEM) to (3) assumes the effects of omitted country-specific variables are fixed and correlate with the regressors. Suppose agglomeration variables compose the vector X and INF is not explicitly introduced into X. It then follows that infrastructure effects are probably present in the fixed effects if they do not vary widely over time.[4] If, however, infrastructure varies over time, then they should appear in the vector X and the fixed effects capture something else. The Random Effects Model (REM) treats the country-specific effects as random variables, which are independent of the regressors. If T (time units) is sufficiently large compared to N (cross section units), the two methods yield similar results. Since the nature of country specific

[4] Mody and Srinivasan (1998) mention that any attempt to use the between (cross sectional) variation implies the possibility of bias because of the correlation between ε_{jt} and X_{jt}. If $\varepsilon_{jt} = \mu_j + v_{jt}$, it is possible that the μ_j term represents an important omitted variable correlated with some elements of the X. The bias then occurs because the influence of this unobserved state effect on FDI may be wrongly attributed to state attributes. Mody and Srinivasan (1998, p. 785) point out infrastructure as an obvious candidate for the bias, since it changes slowly from one year to another: "If the unobserved μ_j reflects general business and operating conditions in the country, it is likely that μ_j and infrastructure will be correlated."

effects is unknown in FDI analysis, we estimate both models and compare the results using statistic-based criteria, such as the Hausman test.

Across all specifications, we use White's matrix of heteroskedasticity consistent covariance, which makes the variance estimator robust to heteroskedasticity. This does not mean, however, that the estimations are free of contemporaneous correlation across cross-sections. In order to check this, we conduct two sorts of serial correlation tests, both derived from the Lagrange Multiplier (LM) Breusch-Godfrey test. For each panel equation we regress the computed residuals on RHS variables and on (lagged one period) residuals. In the annual data context, any serial correlation is likely to appear with only one lagged residual term. The tables below report both the LM t-statistics on the lagged residual term and the LM NR^2 statistic, which has a $\chi^2(p)$ distribution, where p is the number of parameters in the auxiliary regression.

An extension of (3) is to consider a lagged dependent variable model:

$$FDI_{jt} = \alpha_j + \beta X_{jt} + \Sigma \rho_k FDI_{jt-k} + \varepsilon_{jt} \quad (4),$$

upon which first-differencing eliminates the individual effect and produces:

$$\Delta FDI_{jt} = \beta \Delta X_{jt} + \Sigma \rho_k \Delta FDI_{jt-k} + \Delta \varepsilon_{jt} \quad (5),$$

which can be estimated by GMM as suggested by Arellano and Bond (1991) and Blundell and Bond (1998). Efficient GMM techniques employ different types of instruments as discussed, e.g., in the survey by Bond (2002).

THE VARIABLES

We proceed to the construction of independent variables appearing in the vector X: local GDP (Y), industrialization intensity (AGG), "international infrastructure" measured by telephone lines (INF), and government expenditures on infrastructure (GOV). The rationale for choosing these specific variables in the vector X is three-fold. First, INF lies on the foundation of our hypothesis that infrastructure should lead to higher FDI; second, variables other than INF must be present in vector X, upon which expected signs can be assigned; and third, data availability. All real variables are deflated by the Banco de Mexico's 1994 base year Consumer Price Index. The list of variables is as follows:

Foreign Direct Investment (FDI): The dependent variable contains the yearly U.S. dollar amount of foreign investment that each Mexican state receives. Dividing it by population or by state GDP does not change qualitatively the nature of the results.

Market Size (Y): Approximated by the per capita GDP, this variable is constructed by taking the GDP of each state, deflating it by the Banco de Mexico's CPI with base year 1994 and dividing it by each year's state annual population. The annual population is estimated as follows: we take the data given by INEGI in the 1990 demographic census and multiply each year by the INEGI's state average population growth rate of that decade. We expect the larger Y is, the more a state receives of foreign investment due to a larger local demand for the firm's products. This variable, affecting the revenue side of firms, is perhaps the most common in FDI studies, including Coughlin et al. (1991), Braunerhjelm and Svensson (1996), Love and Lage-Hidalgo (2000), Zhang (2001), Deichman et al. (2003), Filippaios et al. (2003), Tuan and Ng (2003), among others.

Infrastructure (INF**)**: The amount of residential and non-residential telephone connections per each thousand state inhabitants. The very same variable has been employed by Bougheas et al. (2000) and Li and Liu (2005) to several countries and by Asiedu (2002) to African economies. The relationship between INF and FDI must be positive, as more firms feel attracted to the state's physical capacity, presumably due to lower operating costs. Three other measures were used: INF1 as the total length of the state's interstate road network (kilometers); INF2 as the total length of the state's paved secondary roads (kilometers); and INF3 as the total square kilometers of the state's paved routes: INF1 + INF2. Theoretical work in Martin and Rogers (1995) distinguish between domestic and international infrastructure. General public administration and transport infrastructure that facilitates domestic trade can be classified as *domestic infrastructure*, while the building of harbors, international airports or the improvement of international communications system are interpreted as *international infrastructure*. If only INF appears to be robust in the estimations, it may indicate that international infrastructure is operative in Mexico. Casual inspection confirms that the Mexican capital and northern border states have larger figures. The INF figures are highest in the DF at 29.11, in Nuevo León at 18.86, and in Baja California at 16.27.

Agglomeration Economies (AGG**)**: Constructed as the share of manufacturing output in a state's GDP. This variable may affect profits in various ways. Coughlin et al. (1991) refer to states with higher densities of manufacturing activity attracting more FDI because foreign investors might be serving existing manufacturers. Firms could also become more efficient due to economies of agglomeration located in the state, since clusters of manufacturing units would attract more firms that wish to profit from such positive external economic effects. The relationship between AGG and FDI must be positive. The AGG figures are highest in Coahuila at 0.35, State of Mexico at 0.33, and Querétaro at 0.32, known for highly concentrated manufacturing clusters.

Government Expenditures on Infrastructure (GOV): The share of the annual state's government expenditures on infrastructure and investment promotion out of its total annual expenditures. We expect that the more a government spends on its state's physical capacity, the more firms feel attracted by these incentives. The relationship should be positive, although of course too large GOV should dismantle the market mechanism and diminish the attractiveness of establishing business in a certain place. Previous studies handling this variable include Deichman et al. (2003), who utilized public investment expenditure as share of provincial GDP.

Real Wages (RW): Two different sorts of real wages are used. First, a weighted tradable and non-tradable sector wage average is taken. The tradable component is the maquiladora industry's real annual average wage in pesos and the non-tradable component is the construction industry's real annual average wage. The tradable weight is estimated as the sum of Manufacturing plus Agricultural and Service Sectors State GDP divided by its total GDP. The non-tradable weight takes into account the rest of the sectors. Second, we simply use the maquiladora's per state annual real average wage in pesos. Over the years, the latter was found to be a parallel shift of the former, and the results are not sensitive to the change in these definitions. While capturing the cost side of profit functions, there is ambiguity in the interpretation of the sign of wages in the FDI literature. This is so because it may represent skill upgrading and therefore might imply a positive correlation between RW and FDI. Because the coefficients were never statistically significant, we remove RW from the benchmark model and keep it as instrumental variable.[5]

Labor Union (LU): This is the ratio of the collective contract demanders over the state's total amount of firms. We expect LU to have a negative relationship with our dependent variable, as more collective demands suggest higher costs. While Beeson and Husted (1989) found that higher levels of

[5] With only two independent variables (Y together with RW), RW have either positive (close to one) or statistically insignificant effects on FDI flows. These results on real wages are not unheard of in the literature. Zhang (2001), for example, finds a –0.027 and statistically insignificant value for a panel of Chinese states, a result also shared by Giulietti et al. (2004) for a panel of UK food sectors. Sun et al. (2002) find for the early subsample wage coefficients of 6.06 with t-value of 2.74 under OLS and 4.66 with a t-value of 6.72 under GLS. They justify the contrary to expected findings in this manner: "Recall that during this early period, most of the FDI originates from Hong Kong… Since the Hong Kong manufacturing industry was export-oriented, goods produced by these Chinese factories were export out eventually. Quality control for these export products was essential. In this period, factory managers with skilled workers were sent to China to train up the local workers. Needless to say, it was quite costly and hence Chinese skilled workers were in big demand." Sun et al. (2002, p. 100). Filippaios et al. (2003, p. 1784) explore U.S. FDI into Australia, New Zealand, Japan, and Korea for 1982-1997 and justify the positive results of wages on FDI similarly: "A possible explanation could be that higher wages is an indicator of a more specialized labour force."

unionization are associated positively with productive efficiency in manufacturing across states, Coughlin et al. (1991) maintains that unionization is expected to deter FDI but do find surprisingly positive effects in his state-level econometric level across the 50 U.S. states. For the U.K. food sector, Giulietti et al. (2004) find that changes in trade union membership have attracted FDI, but only at the 10% level. Though in all panels the expected negative effect of LU on FDI is observed, LU has significance only for the non-border groups. Preliminary estimates suggest LU affects FDI negatively only in states not close to the United States. Exploring in detail this point is an interesting research topic by itself but clearly lies beyond the scope of this paper. Similar to wages, we leave LU to be used as instrumental variable.

THE RESULTS

Analysis of the Benchmark Model

Table 2 contains the REM estimations of the benchmark model with Y, INF, AGG and GOV variables. The maintained hypothesis is exogeneity of these four RHS variables, which will be relaxed later. Hausman tests at the bottom of the table confirm that the REM is well specified since the null hypothesis is never rejected at standard confidence levels.[6] The measure of international infrastructure INF has positive and more than proportional impacts on the FDI flows, although this is weakened for the panel of states of the border. The market size variable is not statistically significant, however. The AGG coefficient is statistically significant with robust values of 0.67 (“all states”) or 0.72 (“all states but DF”). The government size variable has a negative, yet statistically insignificant coefficient, in most cases. The null of no serial correlation is rejected only for the rank panels: -2.48 by the t-statistics for the model with GOV and -2.63 for the model without GOV. With three explanatory variables (output, infrastructure, and agglomeration), similar results on the coefficients are observed.

Following the suggestion by Bekaert and Harvey (1997), we report the Adjusted R^2 statistics of the common intercept model, which is plagued by serial correlation problems. In such model, the constant is the same for all states. For the “all states” panel, for instance, moving from the random effects to the common

[6] Following Betts et al. (2001), we also adopt a compromise strategy between FEM and REM. We estimate REM that control for FDI levels in the U.S. (FDIUS), the neighboring country for Mexico and responsible for a large part of its FDI inflows. The correlation between FDI into Mexico (summing all 22 states) and FDI into the U.S. is quite high. Despite being highly correlated, other regressors than FDIUS appear to capture most of the variation (at least 80%) of FDI into Mexican states. The results of these models are available upon request.

intercept model reduces the Adjusted R^2 statistics from 0.848 to 0.665: a 0.183 reduction. For the "all states but DF" panel, a 0.233 reduction is achieved, while for the "border states" panel there is a 0.288 reduction. State-specific intercepts therefore explain in general about 18% of the variation of flows into the 22 Mexican states covered in the data.

How well do the results of the benchmark model in Table 2 fare in view of the literature? Given the construction of AGG variables, it is recommended we focus on Zhang (2001) and Coughlin et al. (1991). Zhang (2001) verifies AGG as industrialization intensity and obtains a 0.87 coefficient for China, statistically significant at 10%, together with a 0.21 coefficient for the transportation network. Coughlin et al. (1991) also employ manufacturing density and obtain coefficients between 0.355 and 0.473 on the probability of selecting a specific U.S. state for FDI with infrastructure and incentives as explanatory variables.

The coefficient on INF (an index of infrastructure quality) is found to be 1.57 and on AGG is 1.40, both statistically significant in the translog specification on investment by Wheeler and Mody (1992). When lagged FDI captures agglomeration effects, INF (output of electricity per dollar GDP) is found to be below one in general and between 0.67 and 0.74 in the REMs by Mody and Srinivasan (1998). These figures look in agreement with Head et al. (1995): increases in the number of establishments in some industry by 10% would increase its likelihood of being chosen by a subsequent investor in that industry by 5-6%.[7] OLS cross-sections in Asiedu (2002) on developing countries yield coefficients of telephone lines on FDI that range from 0.57 to 0.84, increasing to 1.35 in case of panel data estimation under interactive African country dummy variables.

With only two explanatory variables (output and infrastructure), Table 3 shows that the values of the coefficients associated with real output are higher when infrastructure is measured by roads (INF1) than by phones (INF). The standard errors on the former, however, are very large, yielding lack of statistic significance for the coefficient on INF1. This suggests *that infrastructure has an effect on the volume of FDI only if international infrastructure (telephone lines) is considered*. This finding has been reported by Bougheas et al. (2000) for cross-country studies. Our findings suggest the concentration of telephone lines affects FDI positively at about 2 for the largest panel.

Contrary to the benchmark model, however, the output coefficient is statistically significant for the panel of all states at 1.636. This suggests misspecification when both Y and INF are included in the vector X. We will

[7] A different procedure is adopted by Sun et al. (2002) who calculate the relative accumulation of FDI to domestic investment and find a negative and statistically coefficient of -0.37 on Chinese FDI. They interpret this as existing FDI not attracting further inflows of FDI fast enough, thus bringing a limit to agglomeration (a threshold level).

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return to this point soon. Not reported in Tables 2 and 3 are the estimations of smaller scale models. Results for models with only real state output (Y) as the explanatory variable yield state output coefficients varying from 3.29 for the "rank" group in REM to 5.40 for the same group in FEM. With AGG as the single-variable model, the elasticities of the coefficients in the REM model are reduced. Serial correlation implies badly specified regressions when both Y and AGG are sole explanatory variables.

Model Specification Issues: Residuals

What can be said about the model specification of FDI determinants? Does the move from larger scale vectors to a small-scale vector X of independent variables lead to improvements in the overall fit of the different models? Apparently not, since the estimated effects of regressors on Mexican state FDI become incredibly large in the smaller type models. One strategy is to inspect the residuals of each model since they represent deviations from the predicted FDI values to observed FDI figures. We do not find substantial outliers in the residuals, which in general are well below the critical 1.65 level to be classified as outlier. Apart from occasional values for the Guerrero and Guanajuato states in some specifications, only the residual of the San Luis Potosi state was systematically negative in 1998, suggesting that FDI into that state was lower than predicted by the model for that year.

Table 4 reports the (averaged over time) residuals generated by the REM for the "all states" panel. We calculate the standard deviation of time series residuals for each state to identify whether FDI that can be explained by the model. The standard deviations of residuals are the entries shown in the upper part of Table 4. The bold-faced cells represent the highest (or "worst") residuals across each cross-section unit. The sum of these bold-faced figures across states is reported in the line "Sum of Highest Residuals across All States". The procedure is repeated and summaries appear in the bottom part of Table 4.

It follows from Table 4 that the variability of the residuals is substantially larger under the simplest model with AGG only. For the "all states" panel, averaged residuals over time are more volatile for the AGG model for 13 of the states and for 5 of the states for the largest model (FDI, Y, INF, AGG, GOV). The specification with just AGG in vector X is consistent with infrastructure not changing much over time. According to Table 4, this specification turns out to be the worst of all. The highest residuals for the model with AGG only are also observed redefining panels. For the 21 states panel (removing Mexico City: DF), 12 states present highest residuals in the simplest model with only AGG only and 3 states do so for the largest model.[8]

Overall, the analysis of residuals across the five models suggests that either the largest X vector (the benchmark) or the model with (Y, INF) as explanatory variables are the preferred ones. In any case, INF turns out to appear as a legitimate (RHS) variable rather than appearing as the constant term of panel data estimations. This is consistent with plenty of variation in infrastructure from year to year, which certainly has been the case in Mexico after abolishment of restrictions on foreign capital and NAFTA from 1994 onwards. See also footnote 4 on this methodological note.

Model Specification Issues: Endogeneity

Also critical is the bidirectional effect of output on FDI and of FDI on output. In order to address this point, we check first the sample correlation matrix. We see very high correlation coefficients between INF and Y (0.93), followed by INF and FDI (0.70) and by FDI and Y (0.63). All other correlation coefficients are smaller than 0.40. When infrastructure is measured by roads (in kilometers), the correlation coefficients between INF1 and Y is -0.184 and between INF1 and FDI is -0.33.

To examine whether there exist endogenous relationships between FDI, output and infrastructure we perform augmented Durbin-Wu-Hausman (DWH) tests as done recently by Li and Liu (2005). Let, for example, FDI be expressed as function of Y and INF while Y is written as function of FDI and AGG. We can estimate FDI as function of Y, INF and AGG and generate the residuals, which are then included in an equation of Y estimated on FDI, AGG. If the included residuals are statistically significant, there is endogeneity between FDI and Y. We obtain by the DWH a coefficient of -0.173 (standard error of 0.009), which represents a t-statistic of -19.475. Applying the same procedure for FDI as function of Y, INF, AGG and GOV and Y as function of FDI and RW, the DWH yields a t-statistic of -6.713. It is fair to assume therefore *an endogenous relationship between FDI and Y*.

Applying the same DWH procedure, *we do not find endogeneity between FDI and INF*. If we assume FDI can be expressed as function of Y and INF, with INF modeled as function of FDI and Y, the coefficient associated with the residuals is 0.009 (standard error of 0.057): a t-statistic of only 0.154. Yet there

[8] Table 4 also reports the result of residual checking when classified by lowest values of volatility. For example, for 8 states the residuals have lowest standard deviation in the more general model with (FDI, Y, INF, AGG, GOV) for the "all states" panel; and 7 states do so for the "all but DF" panel. The larger dimension model is thus preferred.

appears that collineality plagues the relationship when Y and INF appear together in X given their high sample correlation.

Under the assumption of *an endogenous relationship between FDI and Y*, one can estimate a dynamic panel data by the generalized method of moments (GMM) as suggested by Arellano and Bond (1991) and Blundell and Bond (1998). Critical in any instrumental variable (IV) procedure such as the GMM is the selection of appropriate instruments. We use RHS variables as instruments and estimate by panel GMM equation (3) above and versions of equations (4) and (5). Efficient GMM techniques employ different instruments as discussed by Bond (2002).

Table 5 contains estimations in levels and first-differences of FDI as function of explanatory variables. Infrastructure captured as international phones has a strong and statistically significant effect on FDI, with the estimates varying from 2.717 to 3.354. These results are not sensitive to instrumenting by either lags of other X's variables such as RW and LU or by lags of AGG and INF. Column (3) with the specification (INF, AGG, GOV) suffers from serial correlation problems. With detailed serial correlation tests omitted for space constraints, the results of Table 5 reinforce the explanatory role of INF in the REM models of Table 2. The set of results under dynamic GMM suggests that INF is the only statistically significant explanatory variable in FDI equations.

The right panel of Table 5 contains estimations in first-differences. Estimations under other instrumental variables were performed and did not change the qualitative conclusions. Increases in infrastructure have a positive effect on changes in FDI with a more than proportional effect. The fit of the first-differenced equations, however, are not adequate as evidenced by serial correlation problems and negative Adjusted R^2 statistics.[9] Despite these, the dynamic GMM estimates confirm the previous findings on INF by taking into account explicitly the ongoing endogeneity between FDI and output.

[9] One explanation for this finding is that parameters may not be identified using first-differenced GMM when the series are random walks as noted by Bond (2002). Another possibility is that we do not have sufficient observations in the time dimension to handle estimations in differenced form. Equations with lagged FDI as regressors suggested considerable persistence in the series as noticed by the close to one coefficients associated with the lagged FDI regressor.

CONCLUDING REMARKS

Despite Mexico's resolute opening to foreign capital from 1994 onwards, there has been no systematic study on the relationship between FDI and infrastructure. We conduct estimations at Mexican states over 1994-2001, following a "general to specific" strategy. The concentration of telephone lines appear to be very important to FDI as their coefficients are about 2.0 in panel REM estimations. Agglomeration is also important, with coefficients varying from 0.62 to 0.67. The article discusses the robustness of these findings along several dimensions, including endogeneity between FDI and real GDP and collineality problems between INF and Y. Under dynamic GMM panels, the effect of telephone lines on FDI increases to around 3.0.

Two main conclusions are offered. First, international infrastructure (telephone lines) appears to be more relevant to Mexican FDI than domestic infrastructure (interstate roads). While Martin and Rogers (1995) provide theoretical discussion on this matter, no empirical evidence seems to exist for FDI. Second, this study argues that treating infrastructure as a constant leads to specification problems. Our findings support ongoing conventional wisdom in Mexico that *investment in infrastructure attracts FDI*.

Due to data constraints, this paper studies aggregate FDI flows into Mexico. Recent studies by Calderón et al. (2004) have emphasized that acquisition of existing assets (M&A) grew much more rapidly than "greenfield" FDI. Distinguishing between the two types of FDI could provide further insights into the role of agglomeration and infrastructure in the process of attracting FDI. Recent works, such as Albuquerque (2003), suggest that FDI should be relatively higher for countries with greater financing constraints. These are two possible research routes for the future.

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Figure 1

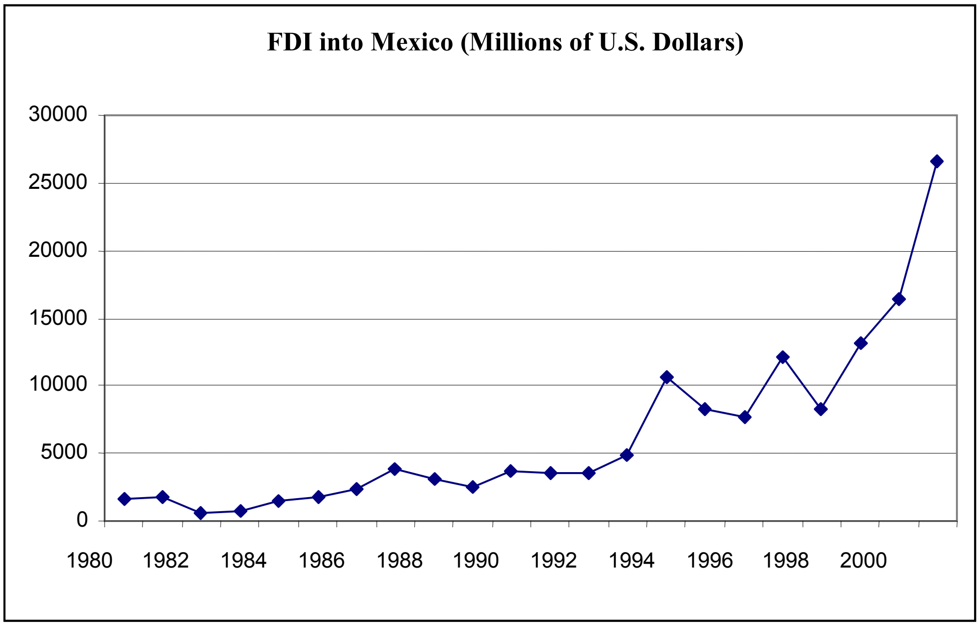
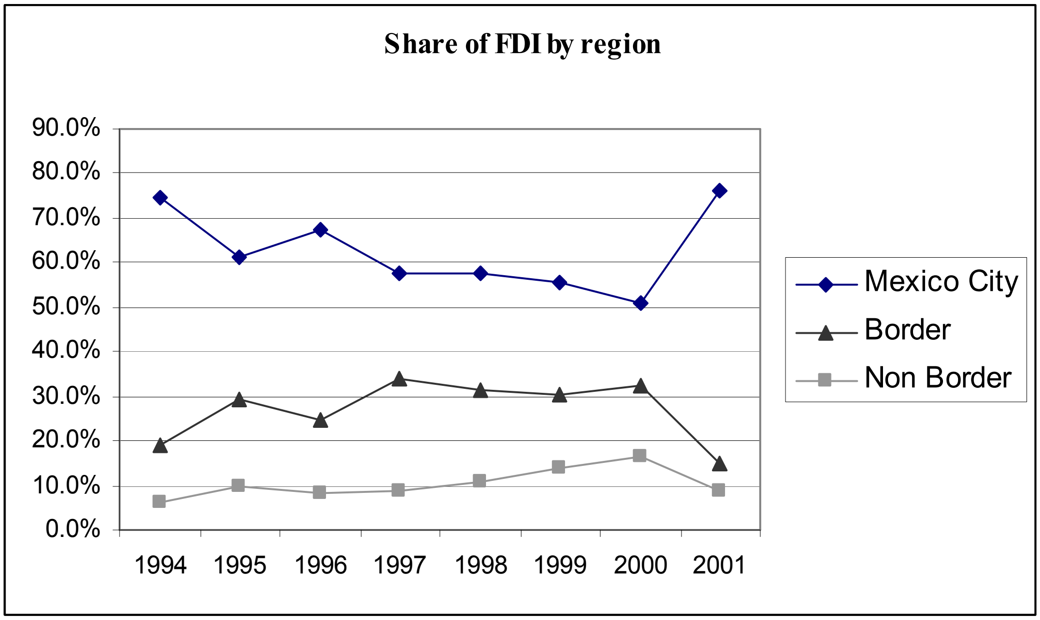

FDI into Mexico (Millions of U.S. Dollars)
30000
25000
20000
15000
10000
5000
0
1980
1982
1984
1986
1988
1990
1992
1994
1996
1998
2000

Figure 2



DOI: 10.2202/1524-5861.1094

Table 1
Flows of FDI into Mexican States in Millions of U.S. Dollars, 1994-2001.

	1994 -1997		1998 -2001		Average per Year	
State	Flows	Share	Flows	Share	Flows	Share
National total	*38872.5*	*100.0%*	*64425.6*	*100.0%*	*12912.3*	*100.0%*
Mexico City	*25251.9*	*65.0%*	*40738.2*	*63.2%*	*8248.8*	*63.9%*
Distrito Federal	23612.9	60.7%	37449.3	58.1%	7632.8	59.1%
Edo. De México	1639.0	4.2%	3288.9	5.1%	616.0	4.8%
U.S. Border States	*10469.6*	*26.9%*	*15911.7*	*24.7%*	*3297.7*	*25.5%*
Nuevo León	4336.5	11.2%	5919.7	9.2%	1282.0	9.9%
Baja California	1870.5	4.8%	2845.4	4.4%	691.2	5.4%
Chihuahua	1875.1	4.8%	2242.7	3.5%	607.5	4.7%
Tamaulipas	1374.1	3.5%	1295.7	2.0%	375.7	2.9%
Sonora	529.2	1.4%	776.5	1.2%	185.2	1.4%
Coahuila	484.2	1.2%	603.7	0.9%	156.1	1.2%
Non-U.S. Border States	*3151*	*8.1%*	*7775.7*	*12.1%*	*1365.8*	*10.6%*
Jalisco	566.2	1.5%	2012.4	3.1%	377.8	2.9%
Puebla	473.5	1.2%	787.4	1.2%	218.7	1.7%
Querétaro	324.4	0.8%	419.9	0.7%	115.2	0.9%
San Luis Potosí	175.6	0.5%	507.1	0.8%	111.2	0.9%
Guanajuato	84.5	0.2%	215.5	0.3%	64.6	0.5%
Baja California Sur	105.3	0.3%	223	0.3%	58.1	0.4%
Morelos	165.6	0.4%	251.9	0.4%	54.5	0.4%
Aguascalientes	108.5	0.3%	207.6	0.3%	50.8	0.4%
Yucatán	129.7	0.3%	127.2	0.2%	49.4	0.4%
Sinaloa	204.7	0.5%	66.7	0.1%	41.4	0.3%
Guerrero	64.3	0.2%	48.9	0.1%	16.7	0.1%
Durango	67.3	0.2%	40.6	0.1%	14.7	0.1%
Tlaxcala	41.7	0.1%	57.8	0.1%	14.1	0.1%
Zacatecas	50.7	0.1%	36.9	0.1%	11.6	0.1%
Other states	589.0	1.5%	2772.8	4.3%	166.9	1.3%

Sources and Notes: Calculations are based on the Banco de Información Económica of Mexico's INEGI. FDI flows for the two periods consider the cumulative amounts of millions of U.S. dollars converted into constant prices. Omitted states are: Campeche, Colima, Chiapas, Hidalgo, Nayarit, Michoacán, Oaxaca, Quintana Roo and Veracruz, which are not included because of data deficiencies. In addition to the regional groups, the empirical analysis includes a group defined as "Rank". It contains the 10 states with the highest average amount of FDI over the sample (1994-2001). This group is composed of the six border states plus the DF, Estado de Mexico, Jalisco and Puebla. Altogether, these 10 states account for about 95% of the total FDI received during the period of analysis.

Table 2
Panel Estimates of FDI Location Determinants
FDI = F1 (Y, INF, AGG, GOV) and FDI = F3 (Y, INF, AGG)
Pooled OLS and FGLS Estimations: Random Effects
1995 – 2001

Panels	All States	All States but D.F.	Border States	Rank	All States	All States but D.F.	Border States	Rank
Y	1.456*	1.199	1.711	0.276	1.351*	1.131	1.838	0.146
Std. Error	*(0.783)*	*(0.855)*	*(1.408)*	*(0.948)*	*(0.771)*	*(0.845)*	*(1.293)*	*(0.910)*
INF	1.856***	1.801***	1.292*	2.582***	2.020***	1.979***	1.278*	2.723***
Std. Error	*(0.586)*	*(0.600)*	*(0.752)*	*(0.728)*	*(0.556)*	*(0.573)*	*(0.710)*	*(0.678)*
AGG	0.666**	0.720**	0.051	0.918	0.621**	0.668**	-0.375	1.022
Std. Error	*(0.310)*	*(0.320)*	*(1.375)*	*(0.931)*	*(0.303)*	*(0.308)*	*(1.123)*	*(0.877)*
GOV	-0.135	-0.154	0.068	-0.087				
Std. Error	*(0.156)*	*(0.163)*	*(0.157)*	*(0.154)*				
Adj. R^2	0.848	0.802	0.853	0.844	0.848	0.801	0.851	0.845
DW	1.656	1.666	2.875	2.181	1.638	1.637	2.787	2.162
Adj. R^2 Com.	0.665	0.569	0.565	0.639	0.665	0.572	0.573	0.639
N	22	21	6	10	22	21	6	10
N*T	153	146	42	70	153	146	42	70
LM t-st.	0.22	0.24	N.A.	-2.48**	0.36	0.40	-1.57	-2.63**
LM NR^2	8.19	7.44	N.A.	5.40	9.75**	8.68	1.26	7.38
Hausman Test	3.448 [0.486]	3.458 [0.484]	0.000 [1.000]	0.000 [1.000]	1.507 [0.68]	2.223 [0.53]	0.000 [1.000]	1.513 [0.68]

Notes: Data are of annual frequency from 1994 to 2001 per state; the 1994 observation is not available for infrastructure, which makes the model run from 1995 to 2001. All estimates are performed by the Feasible Generalized Least Squares (FGLS) method for the Random Effects Model. No weighting is assumed on the residual covariance matrix. The random effects model includes constant terms (β_0) that differ across states. All these constant terms are omitted for space constraints. The "Adj. R^2 Com" refers to the Adj. R^2 of the common intercept model suggested by Bekaert and Harvey (1997) as a test of how much of the variation in FDI is explained by the specified variables. The LM t-stat. is the t-statistic associated with the lagged residual within a standard Lagrange Multiplier test on the residuals of the panel data regression. The LM NR^2 stat. is the value derived from N and R^2 computed in this auxiliary regression. This statistic follows a chi-squared distribution with degrees of freedom equal to the number of estimated parameters (p) in the auxiliary regression: χ^2 (5) = 11.07 for fixed effects and χ^2 (6) = 12.59 for random effects. The LM NR^2 stat. is calculated under the null hypothesis of no serial correlation up to lag order 1. The asterisks *, **, and *** indicate significance levels of 10%, 5% and 1%, respectively.

Table 3
Panel Estimates of FDI Location Determinants
FDI = F4 (Y, INF) and FDI = F4 (Y, INF1)
Pooled OLS and FGLS Estimations: Random Effects
1995 – 2001

Panels	All States	All States but D.F.	Border States	Rank	All States	All States but D.F.	Border States	Rank
Y	4.068***	3.704***	3.415***	2.717***	1.636**	1.553*	1.535	0.101
Std. Error	*(0.470)*	*(0.600)*	*(0.710)*	*(0.550)*	*(0.796)*	*(0.874)*	*(1.068)*	*(0.881)*
INF					2.001***	1.956***	1.408**	2.727***
Std. Error					*(0.563)*	*(0.582)*	*(0.656)*	*(0.672)*
INF1	0.342	0.609	0.380	-0.332				
Std. Error	*(0.210)*	(0.406)	(2.051)	(0.308)				
Adj. R^2	0.261	0.215	0.354	0.316	0.849	0.802	0.852	0.841
DW	1.596	1.591	2.624	1.682	1.632	1.632	2.731	2.107
Adj. R^2 Com.	0.619	0.503	0.556	0.643	0.619	0.503	0.556	0.643
N	22	21	6	10	22	21	6	10
N*T	153	146	42	70	153	146	42	70
LM t-stat.	1.16	1.94	0.221	2.069**	0.33	0.33	-0.52	-2.26**
LM NR^2	0.37	0.81	1.34	1.45	10.92**	9.92**	10.08**	7.08
Hausman Test	3.076 [0.215]	3.905 [0.142]	0.479 [0.787]	11.07*** [0.004]	0.958 [0.619]	1.444 [0.486]	0.000 [1.000]	2.584 [0.274]

Notes: Data are of annual frequency from 1994 to 2001 per state; the 1994 observation is not available for infrastructure, which makes the model run from 1995 to 2001. All estimates are performed by the Feasible Generalized Least Squares (FGLS) method for the Random Effects Model. No weighting is assumed on the residual covariance matrix. The random effects model includes constant terms (β_0) that differ across states. All these constant terms are omitted for space constraints. The "Adj. R^2 Com" refers to the Adj. R^2 of the common intercept model suggested by Bekaert and Harvey (1997) as a test of how much of the variation in FDI is explained by the specified variables. The LM t-stat. is the t-statistic associated with the lagged residual within a standard Lagrange Multiplier test on the residuals of the panel data regression. The LM NR^2 stat. is the value derived from N and R^2 computed in this auxiliary regression. This statistic follows a chi-squared distribution with degrees of freedom equal to the number of estimated parameters (p) in the auxiliary regression: χ^2 (5) = 11.07 for fixed effects and χ^2 (6) = 12.59 for random effects. The LM NR^2 stat. is calculated under the null hypothesis of no serial correlation up to lag order 1, which is reasonable for annual data. The asterisks *, **, and *** indicate significance levels of 10%, 5% and 1%, respectively.

Table 4
Average Standard Deviations of Residuals of FDI Models
Feasible GLS Estimations (Random Effects)

σ of residuals (year average)	F(Y)	F(AGG)	F(Y,INF)	F(Y,INF,AGG)	F(Y,INF,AGG, GOV)	
Aguascalientes	0.418	**0.546**	0.396	0.399	0.414	
Baja Calif.	0.192	**0.284**	0.209	0.210	0.213	
Baja Calif. Sur	**0.529**	0.431	0.367	0.308	0.317	
Coahuila	0.235	**0.287**	0.2215	0.2207	0.2212	
Chihuahua	0.190	**0.282**	0.199	0.169	0.131	
DF	0.488	**0.530**	0.371	0.379	0.365	
Durango	0.838	0.849	1.070	**1.078**	1.046	
Guanajuato	1.389	**1.422**	1.180	1.172	1.159	
Guerrero	1.090	1.101	1.076	1.084	**1.115**	
Jalisco	0.571	**0.751**	0.474	0.479	0.519	
México	0.534	0.511	0.542	0.539	**0.544**	
Morelos	0.750	0.715	0.804	0.823	**0.853**	
Nuevo León	0.590	**0.711**	0.571	0.572	0.570	
Puebla	1.128	**1.276**	1.050	1.043	1.000	
Queretaro	0.189	**0.436**	0.181	0.172	0.164	
San Luis Potosi	1.593	**1.636**	1.456	1.463	1.440	
Sinaloa	0.752	0.758	0.800	**0.803**	0.781	
Sonora	0.290	**0.400**	0.286	0.273	0.279	
Tamaulipas	0.219	0.162	0.287	0.294	**0.344**	
Tlaxcala	0.842	0.793	0.832	0.828	**0.844**	
Yucatán	0.607	**0.671**	0.582	0.573	0.562	
Zacatecas	0.317	0.332	0.657	**0.673**	0.644	
Sum of Highest Residuals						Total of States
All States	1	13	0	3	5	22
All but DF	2	12	1	3	3	21
Border	0	5	1	0	0	6
Rank	0	8	0	1	1	10
Sum of Lowest Residuals						Total of States
All States	4	4	3	3	8	22
All but DF	4	3	4	3	7	21
Border	0	1	2	1	2	6
Rank	2	2	1	2	3	10

Notes: The table reports the (average over time) standard deviation of residuals generated by the FGLS REM for all states. We calculate the standard deviation of time series residuals for each state to identify the amounts of FDI that can be better explained by the REM. The time series standard deviation of residuals, averaged over time, is shown in the cells. The bold-faced cells represent the highest (or worst) residuals across each cross-section unit. The sum of these across states is reported in the line “Sum of Highest Residuals across All States” (22 states). The same procedure is reproduced for the three other panels: “All but DF” (21 states), “Border” (6 states), and “Rank” (top 10 states by average of FDI received over the time period).

Table 5
Panel GMM Estimates of FDI Determinants
FDI = F (X) and Δ(FDI) = G (ΔX)
All States, 1995 – 2001

	Levels	Levels	Levels	Levels	First-Diff. Model	First-Diff. Model	First-Diff. Model	First-Diff. Model
INF	3.201*	2.887***	3.354*	2.717***				
Std. Error	(1.833)	(0.401)	(1.824)	(0.683)				
AGG	1.859	0.721	2.078	0.769				
Std. Error	(3.394)	(0.475)	(2.308)	(0.524)				
GOV			0.717	-0.192				
Std. Error			(1.066)	(0.505)				
Δ(INF)					12.273**			
Std. Error					(4.964)			
Δ(Y)						-21.841*		
Std. Error						(12.628)		
Δ(AGG)							-8.119	
Std. Error							(6.182)	
Δ(GOV)								-9.916
Std. Error								(22.821)
Instr. List	C RW(-1) LU(-1)	C AGG(-1) INF(-1)	C RW(-1) LU(-1)	C INF(-1) AGG(-1)	C RW(-1) RW(-2)	C RW(-1) RW(-2)	C RW(-1) RW(-2)	C RW(-1) RW(-2)
Adj. R^2	0.305	0.293	0.182	0.274	-0.095	-0.462	-0.074	-18.428
DW	1.427	2.013	0.822	2.047	2.954	2.351	2.820	1.911
J-stat.	0.000	0.000	0.000	0.000	0.186	0.042	2.022	0.039

Notes: A constant term is included in all estimations and all series are in logarithms. Data are of annual frequency from 1994 to 2001 per state; the 1994 observation is not available for infrastructure, which makes the model run from 1995 to 2001. All estimates are performed by Efficient GMM dynamic panel methods, with random specification effects, cross-section weights and White (diagonal) for the covariance method. The J-statistic tests the overidentified restrictions E $[Z_i' \Delta\varepsilon_i] = 0$ where Z is the vector of instrumental variables and ε is the error term. It follows a χ^2 (q) distribution where q is the number of overidentified restrictions. At the 5% level, for example, χ^2 (1) = 3.841. The asterisks *, **, and *** indicate significance levels of 10%, 5% and 1%, respectively.

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