Three Essays on the Impact of Openness, FDI and Business Law on Economic Growth

by

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ABSTRACT

The first essay explores the relationship between openness and growth. As, Rodriguez and Rodrik (2000) argue, the relation between openness and growth is still an open question. One of the main problems in the assessment of the effect is the endogeneity of the relation. In order to address this issue, this paper applies the identification through heteroskedasticity methodology to estimate the effect of openness on growth while properly controlling for the effect of growth on openness. The results suggest that openness would have a positive effect on growth, although small. This result stands, despite the equally robust effect from growth to openness.

The second essay investigates the impact of legal differences between countries on the export performance of countries' industries. "High-quality" business law can reduce the cost of external finance by removing informational asymmetries and therefore, benefiting external-finance dependent industries disproportionally. This difference in benefit creates a source of comparative advantage for externally-finance dependent industries located in countries with "high quality" business law. The results indicate that the quality of business law does affect the relative export performance of externally-finance dependent industries. It is also found that the level of financial development also disproportionately benefits externally-finance dependent industries, especially those industries with naturally "small" sized firms.

The third essay examines the impact of FDI on productivity growth. The endogenous nature of FDI and productivity growth presents an obstacle for estimating the impact of one on the other, and lead to biased results from the standard econometric models when use for establishing a causal relationship. However, the endogeneity problem can be overcome by the use of a valid instrument. This paper uses the Chinese government's FDI policy shift of the early 1990s as an instrument for the Chinese FDI in a 2SLS analysis. The results indicate that an increase in the level of FDI as a share of existing capital stock causes an increase in the growth rate of productivity, and that foreign capital is far more productive than domestic capital even after controlling for the province fixed-effects.

Thesis Supervisor: Olivier Blanchard
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"For I am convinced that neither death nor life,
neither angels nor demons,
neither the present nor the future,
nor any powers,
neither height nor depth,
nor anything else in all creation,
will be able to separate us from the love of God
that is in Christ Jesus our Lord."

(Romans 8:38-39)

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I served as a Graduate Resident Tutor at Next House (MIT) for 6 years under Housemasters Bora and Liba Mikic. A special thanks to Professor Mikic and Mrs. Mikic for teaching me how to love and care for undergraduate students and making Next House a home away from home. Their overwhelming love and passion for students were both infectious and inspirational and left an impression that will last a lifetime.

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Once again, is openness good for growth?

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Abstract

Rodriguez and Rodrik (2000) argue that the relation between openness and growth is still an open question. One of the main problems in the assessment of the effect is the endogeneity of the relation. In order to address this issue, this paper applies the identification through heteroskedasticity methodology to estimate the effect of openness on growth while properly controlling for the effect of growth on openness. The results suggest that openness would have a positive effect on growth, although small. This result stands, despite the equally robust effect from growth to openness.

Keywords: Openess; Growth; Heteroskedasticity

1. Introduction

A fundamental question in development and international economics is whether higher trade openness helps to improve economic growth. Even though such a question has received tremendous attention in the literature for more than a century, we still do not have, as a profession, a satisfactory answer. In a recent survey, Rodriguez and Rodrik (2000) provide a critical analysis of the main contributions in the past decade and conclude that "the nature of the relationship between trade policy and economic growth remains very much an open question" (p. 266). However, the authors interpret "the persistent

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interest in this area as reflecting the worry that the existing approaches haven't gotten it quite right" (p. 316).

The present paper offers some insights as to how to exploit differences across countries to address one of the main obstacles to answering the question: the problem of simultaneity, which is pervasive in econometric analysis. What most economists would consider good measures of the degree of openness of a country are, unfortunately, closely linked to the level of income. For example, measuring openness as the ratio between the sum of imports plus exports to GDP clearly is a function of the growth rate of the economy—both the numerator and the denominator are linked to the GDP growth. What this implies is that not even the sign of the bias in the standard OLS regression can be assessed.

Instrumentation via lags or other economic indicators does not offer a valid alternative if, respectively, openness is serially correlated or these other variables affect growth as much as trade. Dollar and Kraay (2003) suggest estimating the regression in first differences and instrumenting the change in openness via lagged values of openness, which appear uncorrelated with other factors influencing changes in growth. Unfortunately, the simultaneity bias can extend over time in the case under consideration. For instance, if part of the growth rate in the future is driven by investment today that requires imported goods, then the degree of openness today depends on future growth rate. Hence, using lag values of openness as instruments does not provide a reliable solution. Alternatively, other economic indicators—such as geographic ones—have been used as instruments (see, for example, Frankel and Romer, 1999). However, these instruments have often been criticized as being correlated with income as much as with trade. For example, the gravity literature has shown that geographic variables can be robustly employed to explain trade. At the same time, they can also be expected to affect growth via other channels, for example, via their relation to health conditions and productivity, to the quality of institutions, and to the availability of natural resources (see Rodriguez and Rodrik, 2000; Baldwin, 2003). To the extent the component of trade that is instrumented via geography is correlated with these other factors, the IV estimates are likely to be biased. Dollar and Kraay (2003) address the instrumentation problem by estimating the growth regressions in first difference and instrumenting the changes in the main explanatory variables (trade and institutions) via their lagged levels. Partly due to the different dynamic nature of the two variables, the lagged levels of trade predict well changes in trade, but not changes in institutions, and vice versa, thus allowing for a better instrumentation for trade.

In this paper, we tackle the extremely important issue of endogeneity again, but from a very different perspective. Instead of appealing to instruments that move the degree of openness but are uncorrelated with income growth (which we have already argued that are difficult to find), we solve the problem of simultaneous equations by using the relatively new literature on "identification through heteroskedasticity" (IH).

In the standard simultaneous equations problem the instrumental variables approach searches for a variable that shifts the demand (for example) to estimate the supply. In other words, we need a variable that moves the *mean* of the demand curve to estimate the supply. We now know that this procedure was introduced in Wright's (1928) book in Appendix B (see Stock and Trebbi, 2003). In the same appendix, Wright indicates that we could also solve the problem of identification if instead of moving the mean we find a variable that increases the variance of one of the equations to infinity. In other words, if the volatility of the innovations to the demand schedule is infinitely large in comparison to the volatility of the innovations of the supply, in the data we only observe movements due to innovations to the demand, and therefore we can estimate the supply schedule directly. This is known today as "near identification".[1] In 1929, Leontief indicated that we did not have to move the variance to infinity, only knowing that it has changed implies that the distribution of the residuals rotate along the different schedules. Rigobon (2003a) shows that this is enough to achieve identification.[2] This approach has been defined in the literature as identification through heteroskedasticity (IH from now on) and has been recently applied to numerous issues plagued by reverse causality.[3]

Sample heteroskedasticity, both across time and across countries, is very high both in country growth rates and in their degrees of openness. It is therefore possible to use the IH method to solve for the problem of endogeneity, when analyzing our fundamental question of the effect of trade openness on growth.

Our results suggest that openness has a small positive effect on growth, which is not particularly robust. They also suggest that most of the empirical works that claim to find a strong and robust result are instead likely to capture either the reverse causality effect or the effect of other economic and policy distortions that are correlated with openness (such as the black market premium, this latter point being already discussed by Rodriguez and Rodrik, 2000).

This paper is organized as follows: Section 2 presents the estimation problem and derives the estimator used in the paper. In Section 3, we present our results using identification through heteroskedasticity. Finally, Section 4 concludes.

2. Endogeneity

As was argued, "good" measures of openness are in general closely related to the growth rate. This generates the standard endogeneity problem. To simplify the discussion, we abstract from other controls and concentrate mainly on the simultaneous equations

[1] See Fisher (1976).

[2] Other theoretical derivations can be found in Sentana (1992) and Sentana and Fiorentini (2001).

[3] Applications where the heteroskedasticity is modeled as a GARCH process are Caporale et al. (2002a), Rigobon (2002a) and Rigobon and Sack (2003a). Applications where the heteroskedasticity is described by regimes shifts are Rigobon (2002b, 2003b), Rigobon and Sack (2003b) and Caporale et al. (2002b). Applications to event study estimation are developed by Rigobon and Sack (2004) and Evans and Lyons (2003). Finally, several applications to panel data can be found in the literature. Hogan and Rigobon (2002) apply the method to a very large panel data to estimate the returns to education. Rigobon and Rodrik (2004) study instead the impact of institutions on income, and how the different types of institutions are affected by income levels and the degree of openness of the country. Klein and Vella (2003) also use heteroskedasticity to estimate the returns to education. Broda and Weinstein (2003) use the inequality constraints together with the heteroskedasticity to estimate the elasticities of substitution in models of trade to evaluate the gains from variety. Pattillo et al. (2003) use the IH method to identify the impact of external debt on growth. Hviding et al. (2004) investigate the impact of official reserves on exchange rate volatility.

problem. Therefore, assume that openness and growth are described by the following system of equations:

$$y_t = \alpha o_t + \varepsilon_t$$
$$o_t = \beta y_t + \eta_t \tag{1}$$

where y is the growth rate, and o is the degree of openness. We are interested in estimating α, but as it is well known that with standard econometric methodologies we cannot consistently estimate α when either β is different from 0 or the variance of η is finite—unless openness and growth are cointegrated—even if we are willing to make the strong assumption that the structural shocks have finite variance and are uncorrelated. The results discussed below apply for any distribution.

Assume we estimate the first equation by OLS not taking into consideration the simultaneous equations problem. The estimation is biased because the right-hand side variables are correlated with the residuals. In particular, in equation:

$$y_t = \alpha o_t + v_t$$

the OLS estimate is given by the general expression:

$$\hat{\alpha} = (o_t' o_t)^{-1} o_t' y_t = \alpha + \beta(1 - \alpha\beta) \frac{\sigma_\varepsilon^2}{\sigma_\eta^2 + \beta^2 \sigma_\varepsilon^2}$$

This equation encompasses all cases where the problem of simultaneous equations disappears:

Exclusion restrictions: if β=0, the bias goes to zero—which should be obvious given that β=0 is assuming the problem away. This is what typically is assumed when exclusion restrictions are imposed on the system of equations. For example, most of the macro literature using VARs and identifying the model using the Choleski decompositions implicitly are making this assumption.

Near identification: Assume that the variance $\sigma\eta^2 \rightarrow \infty$, then it is easy to verify that if the other variance is finite then the bias goes to zero. Near identification is one of the most used assumptions in event study papers. Unfortunately, it is usually assumed implicitly, instead of explicitly.

Cointegration or infinite variance: if the variables are cointegrated, or similarly if the observable variables (y and o) have infinite variance even though the residuals have finite variance, then the expression $(1-\alpha\beta)$ would be close to zero. In this case, it is evident that the bias goes to zero. We know from the cointegrating literature that if the variables are random walks but have a cointegrating relationship (which means that a linear combination of them is stationary), then the model is super consistent and we can run OLS. In this model, if $(1-\alpha\beta)$=0 we have a similar result. Having infinite variance for each of the variables, but finite variance for the linear combination of them is equivalent to having two cointegrating relations: y_t=αo_t+v_t and o_t=β y_t+π_t. The problem in practice is that we will not know which of the two we are estimating, but there will be no inconsistency in the estimates.

There are other assumptions that have been used to achieve identification: long-run restrictions usually assume that the sum of lag coefficients is zero for some type of shock. Additionally, partial identification can be achieved if sign restrictions are imposed. We believe that these assumptions do not apply to the problem of growth and openness, and therefore, we do not discuss them.

Observe that we can also have an estimate for the coefficient in the second equation. The estimate for β is also biased.

$$\hat{\beta} = (y_t' y_t)^{-1} y_t' o_t = \beta + \alpha(1 - \alpha\beta)\frac{\sigma_\eta^2}{\alpha^2\sigma_\eta^2 + \sigma_\varepsilon^2}$$

In summary, both estimates are biased if we cannot justify exclusion restrictions, or that one of the variances is infinitely large, or that the variables are cointegrated, or that there are long-run restrictions (which is a form of exclusion restriction). We believe that the problem of simultaneity between growth and openness cannot be solved by appealing to these assumptions. Indeed, most of the literature does not appeal to them because they are impossible to rationalize in this particular framework.

2.1. Reversed regressions: OLS bounds

Before proceeding towards the IH methodology, it is interesting to discuss a very old literature studying the "bounds" of the OLS estimates in the presence of misspecification from the correlation of the explanatory variables and the residuals (this is a general problem which encompasses other issue in addition to simultaneous equations). The purpose of the bounds is to highlight or show the extent of the misspecification. The method was used by Leontief (1929) and it was recovered by Leamer (1981) and Edwards (1992).

Assume we have a general problem of misspecification that can be summarized by the simple relationship

$$y_t = a o_t v_t$$

where the right-hand side variable o_t is correlated with the residual v_t. Notice that this is exactly the first equation in our system of equations, but here we would like to offer the general discussion when this correlation arises from multiple sources, not just from reverse causality (and hence we use different terms for parameters and residuals than in Eq. (1)).

It is well known that, in the presence of this misspecification, we cannot estimate a consistently. Indeed, there are two forms of estimating a.

$$y_t = a o_t + v_t \tag{2a}$$

$$o_t = \frac{1}{a} y_t + \tilde{v}_t = b y_t + \tilde{v}_t \tag{2b}$$

It is important to indicate that both regressions are *equally* wrong! Leontief studied this problem and realized that depending on the sources of the misspecification the OLS estimates in these two regressions provide bounds for the true coefficient. The estimate in Eq. (2a) provides one bound, and the inverse of the estimate on Eq. (2b) provides the other bound. A special case arises when the misspecification in the model is due to simultaneous equations. In particular, assume output and openness are given by our model Eq. (1). Then the OLS estimate in Eq. (2a) is (the same as before):

$$\hat{a}_{2a} = (o_t' o_t)^{-1} o_t' y_t = \alpha + \beta(1-\alpha\beta)\frac{\sigma_\varepsilon^2}{\sigma_\eta^2 + \beta^2\sigma_\varepsilon^2}$$

while the estimate of $1/a$ in Eq. (2b) is (note that the two expressions are identical):

$$\hat{b}_{2b} = (y_t' y_t)^{-1} y_t' o_t = \beta + \alpha(1-\alpha\beta)\frac{\sigma_\eta^2}{\alpha^2\sigma_\eta^2 + \sigma_\varepsilon^2} = \frac{1}{\alpha} - \frac{1}{\alpha}(1-\alpha\beta)\frac{\sigma_\varepsilon^2}{\alpha^2\sigma_\eta^2 + \sigma_\varepsilon^2}$$

Note that if we are interested in the estimation of α, we want to solve $\hat{b}_{2b}$ for $1/\alpha$ instead of β. We can in fact use both estimates $\hat{a}_{2a}$ and $\hat{b}_{2b}$ to compute the range where the true coefficient α must lie if the model is correct. To illustrate the range, consider the two possible cases, where α and β have different or similar sign.

If α and β have different signs, the bias in Eq. (2a) makes the OLS coefficient smaller (in absolute value) than the true one. In other words,

$$|\hat{a}_{2a}|<|\alpha$$

Similarly, it is also easy to show that in Eq. (2b) the bias is also toward zero. Hence we can write

$$|\hat{b}_{2b}|<\left|\frac{1}{\alpha}\right|$$

Therefore,

$$|\hat{a}_{2a}|<|\alpha|<\left|\frac{1}{\hat{b}_{2b}}\right|.$$

In other words, if the two schedules have different signs, then the true coefficient lies between these two estimates.

The intuition of this result is very simple. First, it is important to realize that Eq. (2a) is the OLS run in one direction, while Eq. (2b) is the OLS regression in the other one. If the schedules have different signs, simultaneous equations will bias the OLS coefficients toward zero—because the OLS coefficient is a linear combination of the two coefficients one positive, and the other negative. Hence the OLS coefficients in both regressions are smaller in absolute terms than the true ones. However, the coefficient in the first equation (Eq. (2a)) attempts to estimate α and the coefficient in the second one (Eq. (2b)) $1/\alpha$. This is what determines the range.

When the two schedules have the same signs the range of coefficients is different. In this case, the bias in the OLS in both Eqs. (2a) and (2b) is away from zero.[4] So, if both

[4] We abstract here from the case where $\alpha*\beta>1$, i.e., when both the observable variables (y and o) as well as the residuals of Eqs. (2a) and (2b) have infinite variance. This case is not very common and might arise in the presence of misspecification due to the omission of variables that are necessary to achieve cointegration.

coefficients are positive the OLS is larger than the true one, and if the coefficients are negative the OLS ones are smaller than the true ones. Nevertheless, this means that in absolute terms,

$$|\alpha| < \min\left\{ |\hat{a}_{2a}|, \frac{1}{|\hat{b}_{2b}|} \right\}$$

Again, this implies a range of coefficients that is admissible. The intuition in this case follows the same reasoning as before, where the difference is due to the fact that in both equations the estimated coefficients are larger than the OLS ones.

For our purposes, if one has a prior that growth and openness positively affect each other, the reasoning above would lead to the expectation that the true coefficients are smaller than the respective OLS bounds. Note also that each of the OLS estimates has a confidence interval. Hence, the exact bounds would need to take into account such interval at the desired significance level.

2.2. *Identification through heteroskedasticity*

In this paper, we appeal to a different methodology: identification through heteroskedasticity. In this section, we derive the basic estimator following closely Rigobon (2003a). Assume we are interested in estimating first the following simultaneous equation system.

$$y_t = \alpha o_t + \varepsilon_t$$

$$o_t = \beta y_t + \eta_t$$

where ε_t and η_t are the structural innovations. The first equation summarizes the growth equation we are interested in estimating. It measures the effect of openness on growth, and the structural residual has the interpretation of being innovations to growth that are independent of all controls and other shocks. The second equation is the openness equation. It describes how growth affects the degree of openness of the economy. The innovations to this equation are interpreted as those changes in the degree of openness that are not explained by the covariates.

Assume that the innovations have mean zero, are uncorrelated, and identically distributed. Additionally, assume that the coefficients are the same across all realizations.

In this model, the only statistic we can compute from the sample is the covariance matrix of the observable variables—i.e., we can compute the variance of growth, the variance of the openness, and their covariance. However, under our assumptions this covariance matrix is explained by four unknowns: α, β, and the variances of ε_t and η_t. This is the standard identification problem in simultaneous equations—there are fewer equations (moments in this case) than the number of unknowns. Algebraically, the covariance matrix of the reduced form is

$$\Omega = \frac{1}{(1-\alpha\beta)^2}\begin{bmatrix} \sigma_\varepsilon^2 + \beta^2\sigma_\eta^2 & \alpha\sigma_\varepsilon^2 + \beta\sigma_\eta^2 \\ & \alpha^2\sigma_\varepsilon^2 + \sigma_\eta^2 \end{bmatrix}$$

where the left-hand side can be estimated in the data and in the right-hand side we have the theoretical moments.

Assume that the data can be split in two sets according to the heteroskedasticity of the residuals, i.e. that the residuals in these two sets have different variances. Remember that in the original model, we have already imposed that the coefficients are the same across all observations. In these two subsamples, we can estimate two variance covariance matrices:

$$\Omega_1 = \frac{1}{(1-\alpha\beta)^2}\begin{bmatrix} \sigma^2_{\varepsilon,1} + \beta^2\sigma^2_{\eta,1} & \alpha\sigma^2_{\varepsilon,1} + \beta\sigma^2_{\eta,1} \\ & \alpha^2\sigma^2_{\varepsilon,1} + \sigma^2_{\eta,1} \end{bmatrix}$$
$$\Omega_2 = \frac{1}{(1-\alpha\beta)^2}\begin{bmatrix} \sigma^2_{\varepsilon,2} + \beta^2\sigma^2_{\eta,2} & \alpha\sigma^2_{\varepsilon,2} + \beta\sigma^2_{\eta,2} \\ & \alpha^2\sigma^2_{\varepsilon,2} + \sigma^2_{\eta,2} \end{bmatrix}$$

this implies that now there are six moments that can be estimated in the sample, which are explained by six coefficients: the two parameters of interest and four variances. Notice that there are as many equations as unknowns. In the standard literature of system of equations, this means that the system satisfies the *order* conditions. To fully solve the problem, then, we have to verify that the six equations are linearly independent—which is known as the *rank* condition.

As is shown in Rigobon (2003a), this requires that the relative variances of the residuals shift across the subsamples:

$$\frac{\sigma_{\varepsilon,1}}{\sigma_{\eta,1}} \neq \frac{\sigma_{\varepsilon,2}}{\sigma_{\eta,2}}$$

The intuition why shifts in the variances achieve identification is quite simple and is closely related to the instrumental variable intuition. Consider the standard demand–supply identification problem. If we are interested in estimating the demand schedule, the IV methodology tells us that we need to find some variable or shock that shifts the supply schedule, so the slope of the demand can be computed. So, the standard intuition searches for something that moves the means.

In the IH methodology, we search for something (like a regime change) that shifts the variance instead of the mean. The different variances in the sample provide enough information to identify the coefficients in both equations (direct effect and reverse causality effect). The shift in the variance rotates the ellipse where the residuals are distributed. That rotation in the ellipse occurs along the schedules we are interested in estimating. Note that the analogy can be brought to the limit in the known case of near identification. Assume that the variance of the supply shocks is infinitely large compared to the demand shocks. In this case, the ellipse enlarges along the demand so much that in the limit it becomes the demand. Therefore, even when we do not know when the supply moves, the likelihood that it is moving is one, hence all the variation observed is along the demand equation. This is known as near identification, and even though there is no movement in the mean, the problem of identification has been solved.

As should be evident from the previous derivation, a crucial assumption is the stability of the parameters. Even though this assumption seems unpalatable in many applications, in cross-sectional panel regressions standard IV methods are implicitly assuming it already.

One advantage of the method we describe is that if there are more than two regimes we can test the overidentifying restrictions.

The identification assumption in the IH methodology is that the data is heteroskedastic (which is easily testable) and that the structural shocks are uncorrelated (this is the maintained assumption). We estimate the model using GMM where the moment conditions are the zero correlation among the identified structural errors. On the other hand, estimation using MLE is more difficult given that it is hard to impose the moment condition—so crucial for identification.

2.3. Adding control variables

It should be obvious that adding additional controls has no impact on the identification problem. Assume that output and openness are described by this system of equations.

$$\begin{pmatrix} 1 & -\alpha \\ -\beta & 1 \end{pmatrix}\begin{pmatrix} y_t \\ o_t \end{pmatrix} = \varphi(L)X_t + \Phi(L)\begin{pmatrix} y_t \\ o_t \end{pmatrix} + \begin{pmatrix} \varepsilon_t \\ \eta_t \end{pmatrix} \quad \text{where } \mathbf{A} \equiv \begin{pmatrix} 1 & -\alpha \\ -\beta & 1 \end{pmatrix} \tag{3}$$

where X are the controls or exogenous variables. In this model, if the variables are not cointegrated (which imposes some constraints on $\Phi(L)$) and we cannot impose exclusion restrictions on the exogenous variables (which imposes constraints on $\phi(L)$ not having a single term equal to zero), the problem of identification cannot be solved with standard methodologies.

The reason is that the reduced form in this model, which is

$$\begin{pmatrix} y_t \\ o_t \end{pmatrix} = \mathbf{A}^{-1}\varphi(L)X_t + \mathbf{A}^{-1}\Phi(L)\begin{pmatrix} y_t \\ o_t \end{pmatrix} + \begin{pmatrix} v_{y,t} \\ v_{o,t} \end{pmatrix} \quad \text{where } \begin{pmatrix} v_{y,t} \\ v_{o,t} \end{pmatrix} = \mathbf{A}^{-1}\begin{pmatrix} \varepsilon_t \\ \eta_t \end{pmatrix} \tag{4}$$

cannot be identified from the data. To illustrate this issue, assume that $\mathbf{R}$ is a positive definite matrix that we use to pre- and post-multiply each of the reduced form coefficients as follows:

$$\begin{pmatrix} y_t \\ o_t \end{pmatrix} = (\mathbf{AR})^{-1}(\mathbf{R}\varphi(L))X_t + (\mathbf{AR})^{-1}(\mathbf{R}\Phi(L))\begin{pmatrix} y_t \\ o_t \end{pmatrix} + \begin{pmatrix} v_{y,t} \\ v_{o,t} \end{pmatrix} \tag{5}$$

The moments from Eqs. (4) and (5) are identical—and therefore, because there are no constraints on $\mathbf{R}$ other than positive definite, there exists a continuum of solutions for A that are consistent with the reduced form moments.

Notice that if we know—for example—that one coefficient of $\phi(L)$ is zero for some lag and some exogenous variable x, then pre-multiplying $\phi(L)$ for an arbitrary matrix $\mathbf{R}$ will violate the restriction. Therefore, if we know that one variable is not included in one of the equations, we can identify the system. This is exactly the intuition of exclusion restrictions—which in this case will imply that variable can be used as an instrument.

One interesting aspect of the reduced form model is that $\mathbf{A}^{-1}\phi(L)$ and $\mathbf{A}^{-1}\Phi(L)$ are consistently estimated by using OLS. In other words, if we were to know what the matrix

A is, then recovering all the coefficients of the structural model (3) would be trivial. In other words, the challenging problem is the estimation of matrix **A**. Notice that from Eq. (4) the reduced form residuals have the exact same properties as the endogenous variables. Hence, the easiest procedure is to use a two-step estimator: First, it is possible to estimate the reduced form model (4)—which is similar to estimating a reduced form VAR—and recover their residuals. Second, those residuals can be used, then, to estimate the contemporaneous matrix **A**.

3. Results

We employ standard growth regression variables in a panel of eight periods of 5 years each, spanning from 1961–1965 to 1996–2000, and about 100 countries. The description of the variables and the corresponding main statistics are presented in Appendix A. In particular, in the IH estimation we use four measure of openness, whose sign is adjusted so that a high value means a more open regime: size of trade (share), a tariff indicator (tarind), import duties (impdut1p), and black market premium (bmp).[5,6]

We perform some preliminary regressions of growth on the control variables and on openness with standard methodologies, such as fixed effect or difference-GMM, to derive some benchmark to assess the importance of properly controlling for endogeneity (the results are presented in Appendix A). The first method addresses the omitted variable bias by adding country specific dummies, but cannot control for endogeneity. The second methodology implicitly accounts for fixed effects by estimating the model in first difference, but also attempts to address the endogeneity issue by instrumenting current variables with previous lags (we adopted the option that allows for all lags to be used). However, if variables are serially correlated, which is expected to be the case for openness and growth, lags are not a very good instrument. The results provide a weak evidence for the effect of openness on growth, as only the black market premium indicator is robustly associated with growth.

We now turn to the derivation of the OLS bounds discussed above and then to the implementation of the IH estimation.

3.1. Reversed regression

Before estimating the IH coefficients, it is instructive to analyze the bounds of the true coefficients following Leontief's reversed regressions. Controlling for fixed effects as well as the other standard variables, we computed the OLS and reversed OLS regressions for

[5] Vamvakidis (2002) included the black market premium among the control variables, rather than as a measure of openness. The regressions we present do not include black market premium in the list of control variables, as in several studies it has been used as a proxy for openness. Introducing it as a control variables in all the regression does not alter the thrust of the results.

[6] Edwards (1992, 1998) argues that the relation between openness and growth should be analyzed with as many measures of openness as possible and he uses nine of them. However, several such measures are unemployable in our setup as they are mainly available as a cross-section.

Table 1
OLS estimates with fixed effects

Measure of openness	OLS Eq. (2a) ($\hat{a}_{2a}$)		OLS Eq. (2b) ($\hat{b}_{2b}$)		Bounds for α	
	Point	*T*-stat	Point	*T*-stat	$\hat{a}_{2a}$	$1/\hat{b}_{2b}$
Share	0.0178	1.97	0.4583	1.97	0.0178	2.1822
Tariff index	0.0553	0.70	0.0245	0.70	0.0553	40.8721
Import duty	−0.0054	−0.19	−0.0263	−0.19	−0.0054	−38.0002
Black market premium	0.2116	5.39	0.3450	5.36	0.2116	2.8987

the four measures of openness: Share, Import duties, Tariff index, and Black market premium.[7] Remember from Section 2.1 that these two OLS estimators determine the bounds where either α or β belong. Indeed, this will be used in the second step to determine the validity of the identification. The OLS results are shown in Table 1; the results for α are of course identical to those in Table A2a in Appendix A.[8]

Notice that the point estimates for the effect of openness on growth (α) are marginally significant on the share variable, highly significant on the black market premium variable, and not significant for the others. Apart for the import duty measure (which is insignificant), they are all positive.

The last two columns of Table 1 show the bounds that the true coefficients have to satisfy.[9] As an example, let us consider the black market premium measure. The coefficients α or β have the same sign. Hence, according to Section 2.1, the true coefficient would need to be smaller than the minimum of the two bounds, which in this case is 0.2116. Hence, if our IH estimator corrects properly for endogeneity, it will need to respect such a condition. The main message of Table 1 is that the bounds are very large, indicating that the endogeneity bias is potentially very large.

3.2. IH estimation: standard setup

In this section, we present the results from estimating the impact of openness on growth using the OLS estimates and the IH methodology.

The procedure is the following:

(1) We estimate Eq. (4) and recover the residuals from the VAR. As was argued before. the residuals share the exact same contemporaneous properties as our variables of interest. Initially, we will not introduce lags of growth and openness, to replicate

[7] The Sachs and Warner measure could not be employed because of the peculiar heteroskedasticity pattern that it would involve: high difference of variance across groups and minimal difference within groups.

[8] Given the presence of strong serial correlation, we run the regression also with lags (even though the literature on growth has generally not included lags). The thrust of the results is unchanged.

[9] Remember that each estimate has a confidence interval, which is not taken into account in the last two columns of Table 3, but would need to be taken into account to derive a precise bound. In our particular case, the estimates have the same sign. Hence, the precise bound would be higher—in absolute terms—than what is reported in the last two columns of Table 3 by the corresponding confidence interval at the desired significance level.

standard growth specifications. We will subsequently add a dynamic structure to account for serial correlation.

(2) We estimate the unconditional covariance matrix for each country and split the data in four groups: high–low variance of openness, and high–low variance of growth, where high and low values are defined with respect to the median. As was argued before, we only need two different covariance matrices to solve the problem of identification. By appealing to four covariance matrices, we have an overidentified system of equations and we can evaluate the validity of the model. In the robustness section, we discuss the implications of the different splits.

(3) Given the four covariance matrices, we compute the contemporaneous coefficients by GMM, where the moment conditions for each regime are:

$$\Omega_i = \frac{1}{(1-\alpha\beta)^2}\begin{bmatrix} \sigma^2_{\varepsilon,i} + \beta^2\sigma^2_{\eta,i} & \alpha\sigma^2_{\varepsilon,i} + \beta\sigma^2_{\eta,i} \\ & \alpha^2\sigma^2_{\varepsilon,i} + \sigma^2_{\eta,i} \end{bmatrix}$$

To compute the standard errors, we use the optimal weighting matrix for the GMM. We use a two-step estimation to compute the optimal weight. The estimation is as follows: From the sample, we compute the covariance matrix from each of the group of countries. This provides 12 moments. These moments have to be explained with 10 parameters: 8 structural shock variances, and the 2 coefficients of interest. GMM reduces the weighted distance between these theoretical moments and the sample ones.

(4) Alternatively, it is possible to specify the GMM by minimizing the identification assumption (that the structural shocks are uncorrelated) in the different regimes.

In Table 2, we present the estimates using the IH methodology. Because the IH methodology can estimate the impact of openness on growth and the impact of growth on openness, we present both coefficients: the impact of openness on growth (α) and the impact of growth on the measurement of openness (β).

Let us analyze first the effect of openness on growth (α), which is the focus of the paper. When comparing the IH results (Table 2) and OLS ones (Table 1), we find that—for the measures that were significant in Table 1—the elimination of the endogeneity bias moves the point estimates in the direction which we would have expected. In fact, as discussed in Section 2.1, for the share and black market premium measures the IH estimates should be smaller than the OLS estimates. Indeed, this is the case. The intuition lies in the fact that both variables affect each other positively and the OLS estimates capture a compounding of both effects.

Table 2
IH estimates with fixed effects

Measure of openness	α		β	
	Point	*T*-stat	Point	*T*-stat
Share	0.00396	1.12	0.47996	5.75
Tariff index	0.07427	1.22	0.00997	0.38
Import duty	−0.00083	−0.05	−0.08367	−1.31
Black market	0.18119	7.12	0.13370	4.28

Table 3
IH estimates with fixed effects for different splits

	1	2	3	4
Share	0.003959	−0.005261	0.019170	0.018884
	1.12	−1.67	5.03	4.95
Tariff index	0.074269	0.053254	0.045819	0.049875
	1.22	0.86	0.65	0.70
Import duty	−0.000828	0.010274	−0.008037	−0.003624
	−0.05	0.67	−0.46	−0.20
Black market	0.181195	0.184601	0.208993	0.210563
	7.12	7.70	7.25	7.34

T-stat below coefficients

Regarding the effect of growth on openness (β) for the same two measures, we find a positive and highly significant effect. These results confirm the prior that part of the positive effect of openness on growth found in the literature should actually be ascribed to the reverse causality—i.e., the one that goes from growth to openness.

The next step is to study how robust these results are to changes in the definitions of the regimes. Here, mostly, we have to estimate the coefficients using different splits. Before showing the results it is important to mention that, as it is argued in Rigobon (2003a), the estimates should be consistent to changes in the windows defining the splits. The intuition is the following: we have said that the system of equations is identified if the true data have heteroskedasticity. In other words, the heteroskedasticity provides additional equations that allow us to solve the problem of identification. Now, misspecifications of the windows (or splits) implies that the covariance matrices estimated in the data are linear combination of the true ones. This is the crucial step—that the misspecified model conforms matrices that are linear combination of the true ones! Hence, if the original system of equations has a solution, then the one from the misspecified splits is a linear combination of the original one. Therefore, the solution is the same as the one from the true system of equations, and there is only a loss in efficiency.

The standard split used in the text assumes that each country belongs to a particular group of variance, where high variance is defined as the variance above median. In a second split, we define high variance as the unconditional moment above the mean variance (this is a very small change in the definition of the windows). In a third split, we look at the different periods of 5 years: we compute the cross-sectional covariance for each 5-year period year and treat each 5-year window as a separate regime.[10] The last split is to use 5-year windows again but now group the window in four distinct groups of high–low variance of each of the two endogenous variables. The results are shown below in Table 3. The first two methods use the different volatilities across countries to create the groups, while the second two methods use the different volatility across time to split the data.

Table 3 shows the estimates for the effect of the black market premium on growth for the four different splits. As can be seen, the estimates are generally close even though the

[10] This corresponds to the methodology employed by Pattillo et al. (2003), with a program kindly provided by Rigobon.

Table 4a
OLS estimates with fixed effects for the first split

Measure of openness	OLS Eq. (2a) ($\hat{a}_{2a}$)		OLS Eq. (2b) ($\hat{b}_{2b}$)		Bounds for α	
	Point	*T*-stat	Point	*T*-stat	$\hat{a}_{2a}$	$1/\hat{b}_{2b}$
Share	0.0298	2.79	0.6317	2.76	0.0298	1.58
Tariff index	0.1037	1.19	0.0462	1.19	0.1037	21.64
Import duty	0.0482	1.43	0.1971	1.43	0.0482	5.07
Black market premium	0.1864	4.58	0.3478	4.55	0.1864	2.87

splits involve very different arrangements of the data. Tariff index and Import duty show significant stability of the coefficient across specifications.

The change in the split made the largest impact on the estimates for the Share variable. For such variable, the first two splits produce insignificant coefficients (although they have opposite signs), while the second two splits produce significant, positive, and similar coefficients. We find that first estimate is not statistically different from the other three (even though some are highly different from zero). However, it is possible to reject the hypothesis that the second coefficient is statistically the same to the third or fourth (at 5% confidence). This suggests a rejection of the model when the share variable is used in the estimation.

There are several possible reasons for this rejection when using share, which mainly pertain to specification issues. First, the omission of lags in the endogenous variables (to the extent these lags belong in the specification), which implies that a common shock is unaccounted for. Second, the relative importance of time-series variation versus cross-sectional variation, given that splits 1 and 2 rely mainly on the first variation and splits 3 and 4 on the second one. Other reasons relate to the existence of non-linearities, of a common shock, or of other endogenous variables unaccounted in the specification. We did not find rejections in the other three variables—which signals to us that there is something peculiar with the Share measure that requires further analysis. The next two sections will further address this issue.

3.3. IH estimation: accounting for serial correlation

As was argued before in the OLS section, there are important serial correlations unaccounted for in the typical growth regression. In this section, we evaluate how the estimates change when we include lags in the specification. This is not a standard growth

Table 4b
IH estimates with fixed effects and lags in first step for the first split

Measure of openness	α		β	
	Point	*T*-stat	Point	*T*-stat
Share	−0.00041	−0.10	0.65265	7.41
Tariff index	0.18950	2.41	0.04326	1.10
Import duty	0.02271	1.02	0.13105	1.52
Black market	0.15881	5.59	0.13847	3.39

Table 4c
IH estimates with fixed effects and lags in first step for the four splits

	1	2	3	4
Share	−0.000412	0.001675	0.028462	0.028051
	−0.10	0.39	5.82	5.71
Tariff index	0.189496	0.152150	0.124494	0.123501
	2.41	1.95	1.47	1.46
Import duty	0.022708	0.018349	0.037903	0.038206
	1.02	0.99	1.66	1.67
Black market	0.158811	0.161419	0.184482	0.184519
	5.59	6.03	6.09	6.10

T-stat below coefficients

specification but one that we were interested in exploring to be sure that the heteroskedasticity is not the result of an unmodeled lag structure.

In Table 4a, we present the bounds for the case in which we allow fixed effects and lags. As can be seen, the same result as before is found. The implied bounds for the true coefficients are extremely large. For each of the openness measures, it goes from less than 0.1 to more than 1.5. This indicates the severity of the endogeneity problem.

In Table 4b we present the results comparable to those in Table 2 with the split 1. The results are similar to those in Table 2, regarding the share and black market premium measure. However, the positive effect of openness as measure by tariff on growth becomes significant and the relation between import duty and growth becomes positive in both directions—although it remains insignificant.

Table 4c reports the results comparable to those in Table 3: the effect of openness on growth, with the four splits.[11] As before, the results are very similar to those in Table 3 for share (different coefficients under the first two and the second two splits) and black market premium (very significant and positive). However, the coefficients of the tariff and import duty indices are now positive, much closer to significance, and stable.

3.4. IH estimation: accounting for serial correlation and cross-sectional variation

The previous two subsections show (Tables 3 and 4c) a puzzling result, which appears to be a rejection of the model when using share. The IH coefficients for the effect of share on openness appear to be different when using splits 1 and 2 or splits 3 and 4, whether lags are present or absent in the specification. This puzzling result arises only when using share and not when using any of the other three measures of openness.

Notice that splits 1 and 2 split the data by country characteristics—i.e. countries with large variance in one variance go to some groups, and so on. Under criteria three and four we use time to split the data, either because every 5-year period is one group (split 3), or because 5-year periods with large variance in one variance go to one group (split 4), and so on. This means that in the splits 3 and 4, the IH estimation will rely mainly on the time-

[11] We estimated the model also using different normalizations and weighting matrices for the GMM. The message from those specifications is almost identical to the one shown here (except, obviously, for the point estimates that change with each normalization). Those estimations are not shown to save and can be provided upon request.

Table 5a
OLS estimates with lags and random effects for the first split

Measure of openness	OLS Eq. (2a) ($\hat{a}_{2a}$)		OLS Eq. (2b) ($\hat{b}_{2b}$)		Bounds for α	
	Point	T-stat	Point	T-stat	$\hat{a}_{2a}$	$1/\hat{b}_{2b}$
Share	0.0459	4.26	0.9345	4.18	0.0459	1.07
Tariff index	0.1145	1.18	0.0412	1.17	0.1145	24.29
Import duty	0.0255	0.77	0.1078	0.77	0.0255	9.27
Black market premium	0.1306	2.86	0.2015	2.86	0.1306	4.96

series heteroskedasticity to estimate the coefficients, while splits 1 and 2 will rely relatively more on cross-country variation.

As the regressions in the previous subsections were based on fixed effect, the main cross-sectional variation was automatically eliminated. This could imply that splits 1 and 2 with fixed effects would not allow the IH estimator to correct properly for endogeneity when the main source of variation is cross-sectional. It turns out that share, or the first stage residuals for share (when not controlling for fixed effect), have much larger cross-sectional variation than time-series variation. This could explain why the problem of different coefficients for the splits 1 and 2 versus 3 and 4 was so pronounced with share. Splits 1 and 2 were unreliable for share as they could not pick its main source of variation, and heteroskedasticity is essential for the IH correction of endogeneity.

In order to avoid this problem, in this section we estimate the model using random effects. As the previous subsection has shown the importance of allowing for lags, we retain them in the specification. In Tables 5a and 5b, we reproduce the OLS bounds and point estimates for the random effects case. As can be seen, the bounds continue to be extremely large. In Table 5c, we present the results for all the different splits. Notice that, in comparison to the previous cases, now all variables, including share, have very stable coefficients. Openness as measured by share and black market premium have a positive and significant coefficient, even though the coefficients are smaller than the OLS estimates presented in Table 5a. Also observe that the tariff index and import duty have positive coefficients, although not significant and they are reasonably close to the results from Table 4c.

In summary, we find that the cross-sectional variation is crucial in understanding the effect of share on growth. However, it is important to mention that random effects allow us to use the cross-sectional variation, but may carry an omitted variable bias, as we are not controlling for country-specific effects. Nevertheless, at least we are sure we know that we

Table 5b
IH estimates with lags and random effects for the first split

Measure of openness	α		β	
	Point	T-stat	Point	T-stat
Share	0.02706	6.83	0.82453	10.22
Tariff index	0.13322	1.65	0.04807	1.64
Import duty	0.01761	1.01	0.05726	0.87
Black market	0.12945	4.75	0.07880	2.19

Table 5c
IH estimates with lags and random effects for the four splits

	1	2	3	4
Share	0.027060	0.027568	0.038623	0.039829
	6.83	6.88	9.34	9.61
Tariff index	0.133220	0.118651	0.081517	0.096128
	1.65	1.48	1.00	1.18
Import duty	0.017613	−0.001302	0.023872	0.023151
	1.01	−0.07	1.24	1.19
Black market	0.129449	0.124598	0.129961	0.126375
	4.75	4.61	4.34	4.24

T-stats below coefficients

are properly controlling for endogeneity and we cannot blame reverse causality (i.e. the effect of growth on openness) for the positive and significant result of share on growth.

Hence, Tables 4c and 5c should be jointly considered our preferred specifications. On the one hand, they both control for the dynamic structure. On the other hand, Table 4c properly controls for country-specific effects, while Table 5c ensures that the IH estimation works properly if we allow it to make use of the cross-sectional variation. When reading both Tables 4c and 5c, we find that openness, as measured by share, black market premium, and, to a lesser extent, a tariff index, has a positive and significant effect on growth. Finally, and equally important, our estimates in Tables 4c and 5c are smaller (or not significantly larger) than those from simple OLS presented in Tables 4a and 5a, respecting the bounds conditions identified in Section 2.1. This is exactly the direction we would have expected if endogeneity is a problem in the data.

4. Conclusions

Academics and policy makers have devoted enormous energy to the question of whether openness is good for growth. Most of the evidence is based either on case studies or on regression analysis. We have learned a great deal throughout the last decades by studying both—however, the question is still open. The main inconveniences are that case studies are always hard to replicate and are affected heavily by country idiosyncrasies, while regression analysis is plagued with endogeneity issues.

It should be clear, now, that instruments to solve the problem of simultaneous equations have been impossible to find in this case. Most of the literature moved toward proxies of openness—such as black market premium—as alternatives, with the unfortunate problem of finding variables that might be correlated with other inefficiencies not necessarily related to the degree of openness. As it is, hence, the problem does not seem to have a solution within the standard econometric methodologies. Furthermore, the best available instrument, distance used by Frankel and Romer (1999), cannot account for the time series variation of the openness variables and should also enter the growth equation directly because it can proxy for quality of institutions, etc. Hence, even the best instrument for openness available in the literature has several limitations.

In this paper, we tackle the same question, using similar data, but resorting to a different procedure: identification through heteroskedasticity. This procedure uses instrumental variables that move the variances instead of the means. In the data, it is the case that the variation on second moments is richer than the variation on means, thus providing scope for using heteroskedasticity to estimate the contemporaneous relationships.

We find that most measures of openness would have a positive effect on growth, even when controlling for the effect from growth to openness. Furthermore, we also show that our estimates are smaller than the OLS estimates—exactly what we would have expected if endogeneity is an issue in the data.

Our results are robust to several specifications when openness is measured by trade over GDP and extremely robust when openness is measured by black market premium. However, as pointed out already by Rodriguez and Rodrik (2000) and Baldwin (2003), among others, black market premium is capturing not only openness but also reflects many other economic and policy distortions. Hence, the focus on the trade aspect of openness may be overstated. In other words, the extreme robustness of black market premium may suggest that it is openness in a broad sense—as part of the overall economic, policy, and institutional environment—that is conducive to growth.[12]

We estimated a very simple model in which other variables—notably all those which are typically available on a cross-section basis—have been excluded. Primarily, we have not included the quality of institutions in the estimation (Rigobon and Rodrik, 2004 study this broader case) which could still potentially explain the positive correlation between openness and growth. Furthermore, we have treated some of the control variables as exogenous when some of them could perhaps be considered endogenous. Future research should extend the current methodology to include those aspects.

From the methodological point of view, this paper shows, once again, that the variation that exists in the data can be used to solve identification issues affected by endogeneity. It can therefore be used to investigate other open questions in the growth literature, especially those related to the impact of policies, as these are likely to be dependent on the level of development and growth of a country. These questions are not only extremely important from the theoretical point of view, but they are crucial policy issues that need our attention.

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[12] Baldwin (2003) suggests that: "One can interpret openness in narrow terms to include only import and export taxes or subsidies as well as explicit nontariff distortions of trade or in varying degrees of broadness to cover such matters as exchange-rate policies, domestic taxes and subsidies, competition and other regulatory policies, education policies, the nature of the legal system, the form of government, and the general nature of institutions and culture."

discussions and comments. This paper should not be reported as representing the views of the IMF or IMF policy. The views expressed in this paper are those of the authors and do not necessarily represent those of the IMF or IMF policy.

Appendix A

We mainly employ the dataset used recently by Vamvakidis (2002) and we complement it with two measures on trade openness from the Economic Freedom Network. The panel encompasses eight periods of 5 years each, spanning from 1961–1965 to 1996–2000, and about 100 countries. The main statistics of all variables are presented in Table A1.

Growth is measured by real GDP growth percapita (ypcg). The set of control variables encompasses initial real GDP per capita (ypc0), the ratio of investment to GDP (iy), the level of inflation (infl), the ratio of M2 to GDP (m2), the growth rate of population (popg), the natural logarithm of the level of education (Ledu, note that in the Table A1 the variable is presented in levels) and the age dependency ratio, i.e. the ratio of dependents to working-age population (age).

We have five openness measures: the ratio of the sum of imports and exports to GDP (share), import duties as a percentage of imports (impdutlp), the average years of openness indicated by the Sachs and Warner index (swyo), the difference between official exchange rate and black market rate (bmp) and a tariff indicator which is the average of revenue from taxes on international trade as a percentage of exports plus imports, the mean tariff rate, and the standard deviation of the tariff rates (tarind). The last two measures are from the Economic Freedom Network dataset. All openness measures are adjusted so that a high value means a more open regime.The appendix also reports the results that are obtained by regressing growth on the control variables and on openness, using standard methodologies such as fixed effects and difference

Table A1
Descriptive statistics

Variable	Observations	Mean	S.D.	Min	Max
ypcg	779	2.03	2.94	−8.49	13.55
ypc0	762	7.51	1.57	4.58	10.74
iy	730	21.64	7.12	3.58	52.43
infl	696	34.50	203.73	−3.01	3357.61
m2	658	33.55	25.14	2.45	180.27
popg	808	1.93	1.12	−0.77	6.25
edu	767	44.00	32.70	1.00	147.83
Age	801	0.77	0.18	0.37	1.21
Share	752	64.36	44.32	5.71	393.75
Bmp	555	7.50	3.48	0.00	10.00
Tarind	521	5.83	2.59	0.00	10.00
Swyo	736	0.42	0.48	0.00	1.00
Impdutlp	439	−10.71	8.53	−47.90	0.00

Table A2a
Growth regressions—OLS-fixed

Share	0.0178				
	(1.64)				
Tarind		0.0553			
		(0.62)			
Swyo			0.8113		
			(2.02)**		
Impdut1p				−0.0054	
				(0.18)	
Bmp					0.2116
					(5.36)***
ypc0	−3.9256	−5.9618	−4.2697	−5.7131	−5.5037
	(6.62)***	(7.38)***	(6.94)***	(6.54)***	(7.33)***
Iy	0.1639	0.1674	0.2032	0.2216	0.1669
	(5.86)***	(5.56)***	(7.72)***	(6.18)***	(5.74)***
Infl	−0.0011	−0.0013	−0.0012	−0.0012	−0.0011
	(2.26)**	(2.80)***	(2.67)***	(2.99)***	(2.68)***
m2	−0.0129	0.0199	0.0104	0.0284	0.0236
	(0.74)	(1.18)	(0.70)	(1.73)*	(1.48)
Popg	−0.0755	0.0685	−0.2970	−0.7784	0.0849
	(0.19)	(0.17)	(0.91)	(1.77)*	(0.24)
Ledu	−0.9522	−1.6403	−1.0918	−0.1224	−1.8478
	(2.92)***	(2.01)**	(3.05)***	(0.15)	(2.91)***
Age	−5.7065	−8.9837	−3.9850	−3.2989	−6.6932
	(2.96)***	(3.28)***	(2.12)**	(1.24)	(3.11)***
Constant	45.9811	68.8750	46.6285	56.2517	62.1187
	(7.99)***	(7.72)***	(7.80)***	(6.32)***	(7.85)***
Observations	476	362	445	265	370
R-squared	0.52	0.59	0.54	0.56	0.60

Robust t-statistics in parentheses.
* Significant at 10%.
** Significant at 5%.
*** Significant at 1%.

GMM (Tables A2a and A2b).[13] The results provide a weak evidence for the effect of openness on growth. Only the black market premium indicator is robustly associated with growth (a positive coefficient represent negative effect of the premium on growth), while the Sachs and Warner indicator is significant only in the fixed effect estimation. Note that the Sachs and Warner index is highly dominated by its black market premium component, and is therefore likely to capture the same effect.

[13] We run the difference-GMM with the one-step robust estimator (the one step is preferred for inference on the coefficients). The tests reject the null of no first order autocorrelation, but do not reject the null of no second order autocorrelation (in the presence of second order autocorrelation the estimates would be biased). The robust option does not deliver the Sargan test for the overidentifying restrictions. When we run the one-step homoskedastic estimator, the Sargan test often rejects the null hypothesis that the overidentifying restrictions are valid. However, when we run the difference-GMM with the two-step estimator (which would partially account for heteroskedasticity but is not reliable for inference on coefficients) or with lags (which would account for serial correlation but is not standard in growth literature), the Sargan test accepts the null hypothesis that the overidentifying restrictions are valid. We do not report such additional results, as heteroskedasticity and—to a smaller extent—serial correlation are discussed carefully in the paper.

Table A2b
Growth regressions—difference-GMM

LD.ypcgmm5	0.6192	0.5843	0.6153	0.4976	0.6130
	(9.58)***	(8.69)***	(8.26)***	(5.70)***	(9.08)***
D.share	0.0386				
	(1.54)				
D.tarind		0.1117			
		(0.89)			
D.swyo			0.9671		
			(1.28)		
D.impdut1p				0.0137	
				(0.34)	
D.bmp					0.1448
					(2.77)***
D.iy	0.1143	0.1836	0.1777	0.2396	0.1484
	(2.76)***	(3.92)***	(4.71)***	(4.11)***	(3.50)***
D.infl	−0.0014	−0.0005	−0.0009	−0.0015	−0.0009
	(1.79)*	(0.55)	(1.22)	(1.95)*	(1.32)
D.m2	−0.0092	0.0394	0.0488	0.0350	0.0410
	(0.24)	(0.96)	(1.26)	(1.23)	(1.06)
D.popg	−0.3421	−0.3495	−0.3362	−1.2418	−0.0733
	(0.95)	(0.70)	(1.07)	(2.76)***	(0.18)
D.Ledu	−3.2490	−3.7085	−2.8249	−1.9806	−2.6286
	(4.31)***	(3.62)***	(3.94)***	(2.37)**	(3.42)***
D.age	(1.60)	−6.4731	−8.1916	−1.6137	−8.0952
		(1.38)	(2.11)**	(0.38)	(2.06)**
Constant	0.6563	0.6483	0.3080	0.7743	0.3398
	(2.74)***	(2.52)**	(1.16)	(3.01)***	(1.43)
Observations	312	240	291	174	249
Number of NO	70	67	65	51	66

Robust z-statistics in parentheses.
* Significant at 10%.
** Significant at 5%.
*** Significant at 1%.

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Business Law As a Source of Comparative Advantage

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BUSINESS LAW AS A SOURCE OF COMPARATIVE ADVANTAGE

I. INTRODUCTION

A variety of outcome measures have been used in the law and finance literature to assess the impact of differences between countries' legal regimes. These outcome measures include a country's level of financial development, such as the size of the country's stock market or the extent of private credit available in the economy, the premium for control blocks and firm valuation. This paper adds to this important literature by examining the impact of legal differences between countries based on the export performance of countries' industries. More specifically, can the quality of a country's business law, broadly defined, be a source of comparative advantage for that country's external finance dependent industries?

This paper begins with the assumption, well-motivated by a number of studies, that "high-quality" business law can reduce the cost of external finance in a country. "High-quality" business law may make it more difficult and costly, in a variety of ways, for controlling shareholders and managers to engage in transferring value from external investors to themselves. Moreover, "high-quality" business regulation may remove informational asymmetries that can drive up the cost of external finance. A lower cost of external finance should disproportionately benefit external-finance dependent industries. This disproportionate benefit should create a source of comparative advantage in externally-finance dependent industries for countries with "high quality" business law relative to countries with lower quality business law. Importantly, "high-quality" business law should not create a source of comparative advantage for industries not reliant on external finance if the primary effect of "high-quality" business law is to reduce the cost of external finance.

Based on this reasoning, we test whether external finance dependent industries' export share increases relative to that of industries not reliant on external finance as the quality of a country's business law increases. Our sample consists of fifty-six countries with country import and export figures broken down by thirty-six industries. These thirty-six industries have a wide range of external finance dependency. We find, with economic and statistical significance, that the quality of business law does affect the relative export performance of externally-finance dependent industries. Specifically, we find that the quality of a country's disclosure regime appears to be a particularly important source of comparative advantage for externally-finance dependent industries based on export data.

We also test a number of additional, related hypotheses. Several recent papers have indicated that a country's level of financial development disproportionately benefits external-finance dependent industries. Manova (2006). Consistent with these papers, we also document, using a variety of proxies for financial development, that financial development disproportionately benefits (again in terms of their export performance) externally-finance dependent industries. We also document that externally-finance dependent industries with naturally "small" sized firms are disproportionately benefited by increased levels of financial development. Interestingly, externally-finance dependent industries with naturally "small" sized firms do not appear to be especially benefited relative to industries with "larger" sized firms from increased business law quality.

Finally, we investigate the effect of a liberalized capital account. We hypothesize that a liberalized capital account should disproportionately benefit external finance dependent industries with "larger" sized firms as it is these firms that will be able to raise external finance on the international capital markets. We find strong results that this is, in fact, the case.

Part II will briefly describe the existing literature. Part III will then explain in more detail our hypotheses, and the motivation for them, while Part IV will describe our data. Finally, Part V will present our results.

II. EXISTING LITERATURE

This paper is related to two literatures: the law and finance literature, and the nascent trade and finance literature. There is a substantial law and finance literature that investigates the effect that the differences between countries in their legal regimes have on a variety of outcome measures, such as financial development. As a result of the focus on legal differences, a number of indexes have been developed that attempt to quantify differences in countries' legal regimes along several dimensions. There are indexes that measure, for example, differences in the extent to which countries (or their primary exchange) impose mandatory disclosure requirements on firms, see La Porta, Lopez-de-Silanes, & Shleifer (2006), and differences in the extent to which creditors' interests are legally protected. La Porta, Lopez-de-Silanes, Shleifer & Vishny (1998). These studies have tended to find that "law matters" in the sense that a number of legal differences are correlated with differences across countries in their level of financial development and other outcome measures. In order to control for reverse causation, these studies typically use the "legal origins" of a country as an instrument reasoning that the legal origins of a country is a pre-determined variable.

One important area of research in the law and finance area (as well as in the corporate finance literature) has been on the implications of firms and industries varying in their external finance dependency. External finance dependency is the extent to which firms need to raise capital to fund investment opportunities from external investors, such as issuing equity. In an important paper, Rajan and Zingales (1998) found that the more the median firm in an industry relies on external financing in the United States, the faster firms in this industry grow in countries with good accounting standards relative to firms in those same industries in countries with weaker accounting standards. The Rajan and Zingales measure of the external finance dependency of an industry has been widely used in a number of subsequent studies (see, e.g., Fisman 2003). In related work, Carlin and Mayer (1998) found that industries that substantially rely on issuing equity to finance investment grow faster in countries with better accounting standards.

Beck (2003) showed that countries with higher levels of financial development exported more goods in industries reliant on external financing relative to exports of those same industries in countries with lower levels of financial development. The same basic results were also documented in Becker and Greenberg (2005). Building on this work at the intersection of trade and finance, Manova (2006) also finds that countries with better developed financial systems tend to export relatively more in highly external capital dependent industries and in sectors with fewer collateralizable assets. She also found that equity market liberalizations increase exports disproportionately for sectors more reliant on external finance.

III. Business Law and Comparative Advantage

Our starting point is the hypothesis that a county with "high quality" business law will have a comparative advantage in industries that rely on external finance relative to those same industries located in countries with "low quality" business law. The more externally-finance dependent an industry is, holding all else constant, the greater the comparative advantage for the country with the "high quality" business law will be. In contrast, a higher quality for a country's business law will not translate into a greater comparative advantage for industries not reliant on external finance, or at least to the same degree as for external-finance dependent industries.

Our hypothesis is motivated by two observations. First, if firms in industries that are externally-finance dependent, i.e. need to raise capital to fund investment opportunities from external investors, enjoy a lower cost of external finance in a particular country, than that country should have a comparative advantage in those industries relative to firms in other countries in those same industries which face a higher cost of external finance. The fact that a country has a comparative advantage in externally-finance dependent industries should manifest itself in the industrial composition of the country's exports. The second observation connects the cost of external finance in a country with the quality of that country's business law. There are theoretical and empirical reasons to believe that high-quality business law can enhance the availability and reduce the cost of external finance within that country. Both observations will be briefly discussed as they form the motivation for our hypothesis.

A. Lower Cost of External Finance as a Source of Comparative Advantage

In the standard Herckscher-Ohlin factor model of trade, countries that are relatively "abundant" in a particular factor of production, such as labor, will tend to export goods that use that factor intensively. Put slightly differently, countries have a comparative advantage in whichever factors of production are "abundant" in that country relative to the world endowment of those factors. In line with such a model, one can think of external finance as a factor of production that is more "abundant" in countries where the cost of external finance is comparatively low.[1] As a result, countries where the cost of external finance is comparatively low will tend to export goods that use this factor intensively, i.e., the products of externally-finance dependent industries.[2] This is a potentially important source of comparative advantage as industries differ substantially in terms of their need for external finance (as proxied by the median firm in that industry) and, moreover, countries vary substantially in the quality of their business law.

B. Business Law and the Cost of External Finance

There are a number of reasons why high-quality business law, broadly defined to encompass corporate law, securities regulation and creditor-debtor/bankruptcy law, might reduce the cost and enhance the accessibility of external capital for firms. First, high-quality business law could provide increased assurance that the funds being raised will be used for the highest valued purpose. Prohibitions on self-dealing, improved shareholder rights, increased transparency and other kinds of business regulation could potentially have this effect by reducing the agency costs created by separation of ownership and control and conflicts between controlling and minority shareholders' interests. For residual claimants, improved allocation of capital will improve the value of their claims and, hence, the amount of capital a firm receives for issuing those claims in the first place. There is, in fact, growing empirical evidence that strong business protections can improve a country's allocation of capital (see, e.g., Wurgler 2000). In addition, improved capital allocation might reduce the need of external investors to expend resources monitoring management and controlling shareholders to ensure funds are properly allocated and spent (Lombardo and Pagano 2002).

On a related note, business law can enhance the "pledgeability" of the firm's assets (and the cash-flows generated by the firm's assets) and thereby increase the

[1] A lower cost of external finance in a country, holding all else constant, implies a larger stock of external capital in that country.

[2] Two theoretical papers have focused on financial development as a source of comparative advantage. Baldwin (1989) developed a model in which financial development served as a source of comparative advantage due to an enhanced ability to diversify risk. The lower cost of bearing risk, given the ability to diversify, reduces the cost of production in risky goods in financially developed countries. Kletzer and Bardhan (1987) demonstrate that a country's financial development can be a source of comparative advantage.

availability of external finance and reduce its cost. (Shleifer and Wolfenzon 2002). Consider, as an example, countries with concentrated ownership of firms, such as those in Continental Europe and most of Asia. There is the widespread concern that firm assets might be "tunneled" to other entities, owned by the controlling shareholder of the firm, at below-market prices. This is not an idle concern (see, e.g., Betrand, Mehta and Mullainathan 2002). The possibility of firm assets being transferred to other entities at below-market prices would impede the ability of these assets (and the cash-flows they produce) to provide assurance to external investors that they will receive payment given that the tunneled assets are no longer owned by the firm. Strong fiduciary obligations, mandated disclosure requirements, minority shareholder rights and other types of business law might reduce the incidence and size of such transfers (and, importantly, any deadweight costs associated with such transfers). In the context of debt, strong creditor rights can increase the "pledgeability" of firm assets if strong creditor rights are defined as the ability to seize or control firm assets, in a cost-effective manner, in the event of default.

Mandatory disclosure requirements (and enforcement thereof) could reduce the cost of external finance through two additional mechanisms. First, mandatory disclosure requirements could reduce the well-documented adverse selection cost associated with raising external finance due to asymmetric information (see, e.g., Huang & Stoll 1997). Reducing adverse selection by reducing asymmetric information is value-enhancing as adverse selection makes it less likely that firms with high-valued projects will raise external finance to fund attractive investment opportunities as their shares will sell at a discount to their true value. Second, mandatory disclosure requirements might reduce the level of private information held by traders and, hence, the cost of trading for uninformed investors. If the cost of external capital is set by uninformed investors, reducing private information should reduce the cost of external capital as uninformed investors' expected loses from trading should be lower (Easley, D. and M. O'Hara (2004)).

If high-quality business law is particularly important in reducing the cost of external finance, and there is substantial evidence that business law can have these desirable effects (see generally Ferrell 2005 and papers cited therein), then the presence of such regulation should primarily benefit firms in industries that rely on external finance. The cost of external finance is simply not relevant for firms that do not rely on external financing.

Of course, it might be the case that business law could reduce the overall cost of capital generally and not the cost of external finance *per se*. One would then expect a county with high-quality business law to have a comparative advantage in capital-intensive industries as opposed to just external-finance intensive industries. In this scenario, one would not necessarily expect to see a differential impact of business law on external-finance dependent industries versus non-external finance dependent industries.

C. Additional Hypotheses: Firm Size and Capital Account Liberalization

In addition to the hypothesis that business law can serve as a source of comparative advantage for external-finance dependent industries, we are also interested in exploring the effects on the export performance of industries of two additional factors: the median size of a firm in an industry and a country having a liberalized capital account. Specifically, there are three sets of questions we wish to investigate.

First, does a country's level of financial development disproportionately benefit (in terms of export performance) industries where firm size tends to be naturally "small," holding constant the external finance dependency of that industry? The motivation for this hypothesis is the supposition that larger firms have an easier time raising external capital than smaller firms in less financially developed markets and, hence, small firms will be disproportionately benefited by higher levels of financial development.

There is, in fact, substantial support for the view that small firms often have more difficulty raising external finance in less financially developed markets. For instance, one study found that small firms in the less financially developed regions of Italy have a more difficult time raising external finance than small firms in more financially developed regions of Italy. In contrast, large firms' access to external capital was not affected by their location within Italy (Guiso, Sapienza & Zingales 2004). On a related note, large diversified business groups tend to be the dominant economic actors in less financially developed countries due to their ability to fund investment projects through internally generated funds rather than having to raise those funds from external investors

(Khana & Palepu 2000). The fact that a potentially costly substitute for external finance, diversified business groups, has arisen in less developed financially countries (but no where near to the same degree in financially developed countries) suggests that small firms in less financially developed countries face nontrivial barriers to external finance. Finally, based on a firm survey, Beck, Demirguc-Kent and Maksimovic (2005) report that smaller firms face binding financing constraints more often than larger firms. The results from this survey suggest that this difference between small and large firms is especially acute in less financially developed countries.

Second, does a liberalized capital account disproportionately benefit (again, in terms of export performance) industries where firm size is naturally "large", holding constant the external finance dependency of that industry? Countries with liberalized capital accounts are countries that permit the free flow of capital, such as foreign direct investment and portfolio flows, into and out of a country. The motivation for this hypothesis is that for sufficiently small firms the ability to access the international capital markets as a result of capital account liberalization is unlikely to be an important new source of external capital. It simply might not be in foreign investors' self-interest to take the time to investigate and monitor "small" firms in a foreign country just in order to make a small investment. In contrast, for "larger" firms access to the international capital markets could be an important new source of funding. Informational asymmetries between foreign investors and domestic firms might be less acute for larger firms. There is some evidence that foreign portfolio flows are invested (beyond their market weights) in larger firms (see, e.g., Liljeblom, E. and A. Loflund (2000)). Whether this hypothesis holds true turns, in part, on whether 'larger" firms would, but for capital account liberalization, face some financing constraints in their domestic markets in terms of the cost and availability of the external finance they need. Particularly large firms, which might be more likely to have political connections or well-established domestic reputations, might not be financially constrained even when the international capital markets are closed. Laeven (2003).

Third, does a country's quality of business law disproportionately benefit industries where the firm size is "small," holding constant the external finance dependency of that industry? If there were economies of scale, for example, in privately credibly committing to better governance or higher-quality disclosures then small firms might be at a disadvantage in the absence of law. Reputational markets might work better for large firms that repeatedly raise capital on the market. On the other hand, if business law equally benefits "large" and "small" firms in terms of their cost of external finance (even if financial development perhaps does not), then one would not see a differential effect.

IV. DATA

A. Data Description

Trade Data

The import and export data for fifty-six countries, broken down by the 36 ISIC Revision 2 industry classification, was obtained from the United Nation's Comptrade Database (which has been deflated by World Bank export and price indices). This database provides the average of the exports or imports of a country for an industry over the 1980-1989 period. Our "export share" variable is the export industry-level data by country divided by that country's GDP. Our "trade share" variable is the net exports (exports minus imports) of an industry for a country divided by that country's GDP for the 1980-1989 period.

Quality of a County 뭩Business Law Proxies

The eight proxies for the quality of a country's business law are taken from the law and finance literature. The "anti-self-dealing" index, which measures the degree to which a country provides minority shareholders protection against self-dealing by controlling shareholders, and the updated version of the "anti-directors" index, which measures the degree to which shareholders have rights vis-a-vis management, was taken from Djankov, LaPorta, Lopez-de-Silanes, Shleifer (2006). The "anti-self-dealing" index's two components, the average of the two of which constitutes the "anti-self-dealing" index, were also used. These two components are measures of the extent to which a country's business law empowers the ex-ante private control of self-dealing, defined as the various approval requirements necessary prior to engaging in a self-interested transaction and the immediate disclosures required in connection with undertaking such a transaction, and the ex-post private control of self-dealing, defined as the ease with which minority shareholders can prove wrongdoing as a result of a self-interested transaction. The "public enforcement" variable was also taken from Djankov, LaPorta, Lopez-de-Silanes, Shleifer (2006) which measures the potential fines and criminal penalties imposed by a country as a result of engaging in self-dealing transactions.

Another important proxy for the quality of a country's business law is the quality of a country's mandatory disclosure requirements. The "disclosure" index for countries is taken from La Porta, Lopez-de-Silanes, & Shleifer (2006). The "disclosure" index measures whether a country requires disclosure of such items as the equity ownership structure of publicly-traded firms, compensation of directors and key officers, contracts outside the ordinary course of business, and transactions between the company and its directors, officers or large shareholders. The accounting standards index, which measures the quality of a country's accounting standards, is the International Financial Analysis and Research's accounting index. This index measures, as of 1990, the extent to which firms in a country disclosed in their annual reports ninety potentially important pieces of information. Finally, the creditors' rights index, which measures the legal rights creditors are afforded, was taken from La Porta, Lopez-de-Silanes, Shleifer & Vishny (1998).

The legal origins of a country's legal regime, typically used as an instrument for the quality of a country's business law, is taken from La Porta, Lopez-de-Silanes, Shleifer and Vishny (1997). The legal origins of a country can be British, French, German or Scandinavian.

Capital Account Liberalization

We measure the degree to which a country has a liberalized capital account based on Quinn (1997). Quinn reports, with zero being a totally closed capital account and a score of four being a completely liberalized capital account, a country's capital account liberalization score as of 1958. Quinn also documents the change in a country's capital account liberalization score between 1958 to 1988. From these two pieces of information we calculate a country's capital account liberalization score as of 1988.

Proxies for Industry External Finance Dependency

The external finance dependency of an industry measure is an attempt to quantify the extent to which firms in a particular industry need to raise capital from external investors. The definition of the external finance dependency of an industry, now widely used in the literature, is taken from Rajan and Zingales (1998). External finance dependency of an industry is the median value in that industry of firms' capital expenditures minus cash flow from operations divided by capital expenditures. We will use three different samples to measure the extent to which an industry relies on external financing. External financial dependency of 36 industry classifications based on ISIC Revision 2 is based on U.S. data for all firms in the 1980s; U.S. data for "young" firms (firms that have gone public within the last ten years) in these industries for the 1980s; and, as a robustness check, all U.S. firms in the 1970s.[3] Focusing on "young" firms is potentially informative as it is typically young firms in an industry that tend to rely on external financing. Moreover, it is possible that firms in industries in the United States tends to be more mature and, hence, focusing on young firms in an industry might provide a more useful measure of the need for external finance of firms in other countries with potentially less mature industries.

An example of an industry with a very low score for external finance dependency is the Tobacco industry. Firms in the Tobacco industry tend to generate substantial cash-flows relative to investment opportunities and, as a result, rarely have to raise capital from external investors, such as issuing net additional shares on the capital markets. The Drugs and Pharmaceuticals industry, in sharp contrast, is an industry that has a substantial need to raise capital from external investors and, hence, has a high score for the external finance dependency variable.

Firm Size of an Industry

We define the firm size of an industry as hundred minus the industry's share of employment by firms with less than 500 employees, and, alternatively, the industry share

[3] Manova (2006) reports that the correlation between measures of industry external finance dependency based on 1980s data and 1990s is very high.

of employment by firms with less than 100 employees, less than 20 employees, less than 10 employees and, finally, less than 5 employees (hereinafter 500, 100, 20, 10 and 5 measures of firm size). Therefore, as the "firm size" variable increases, the firm size in the industry increases as a higher "firm size" value represents a larger percentage of employees in the industry not working in either firms with 500, 100, 20, 10 or 5 or less employees depending on which cut-off is used. These industry-share of employment figures are reported in Beck, T. Demirguc-Kent A., Laeven L., and R. Levine (2004), which in turn took these figures from the 1992 U.S. Census data. We use their data, which covers the 36 ISIC Revision 2 industries, in calculating the firm size of an industry. The U.S. Census did not collect industry-share employment figures prior to 1992.

Financial Development Proxies

We use several widely used proxies for a country's financial development.[4] One variable is "Private Credit" which is defined as the value of credits by financial intermediaries to the private sector divided by GDP. We also use, as in Beck (2003), two related variables as a robustness check: the variable "Liquid Liabilities" which is the liquidity liabilities of a country's financial system and the variable "Commercial-Central Bank" which is defined as the ratio of commercial banks' domestic assets divided by commercial and central banks' domestic assets. The "Private Credit" variable is a particularly important and widely used proxy for financial development in the law and finance literature (see, e.g., Manova (2006); Levine, Loayza, and Beck 2000; King & Levine 1993).

"Market Capitalization," another widely used proxy for the level of a country's financial development, is also used. "Market Capitalization" is the value of listed domestic shares on the country's domestic exchanges divided GDP. Two related variables, "Value Traded" and "Turnover", are also used as a robustness check. "Value Traded" equals the value of shares traded on a country's domestic exchanges divided by GDP, while "Turnover" is the value of domestic exchange trades divided by the value of listed domestic shares (Levine and Zervos 1998).

[4] We obtained the financial development data from Thornstein Beck.

Finally, the variable "Total Capitalization" is the sum of the "Private Credit" and "Market Capitalization" variables. As a result, the "Total Capitalization" variable captures financial development as reflected in both the size of the equity markets and the level of financial institutions' credit provision activity. As a result, the "Total Capitalization" variable is arguably the most comprehensive of the financial development measures.

B. Correlations

The correlation matrix for the financial development measures – Private Credit, Liquid Liabilities, Commercial Central Bank, Market Capitalization, Value Traded, Turnover Ratio, and Total Capitalization – is presented in Table I. There is substantial correlation between the different financial development measures. The "Total Capitalization" measure's correlation with the other financial development measures varies from a low of .54, with the Turnover Ratio, to a high of .95, with Private Credit (the latter not being surprising as "Private Credit" is a component of "Total Capitalization").

The correlations between the different proxies for quality of business law is reported in Table II. As is obvious, the correlation between different proxies for business law quality, while often substantial, also varies substantially suggesting that using different indexes for business law quality is potentially informative. Among the highest correlations is between the two indexes that attempt to measure the quality of a country's disclosures: the disclosure index and the accounting standards index with a correlation of .66. The capital account liberalization index, on the other hand, has a very low correlation with the eight indexes for business law quality.

V. Results

A. Financial Development and Comparative Advantage

As previously discussed, several papers, starting with Beck (2003) and most recently Manova (2006), have examined whether financial development can act as a source of comparative advantage. Given this literature, we start our empirical analysis by seeing whether, consistent with these papers, we find that a country's level of financial development can serve as a source of comparative advantage in externally finance dependent industries. Accordingly, we run the following regression:

$$\text{Exp}_{ic} = \alpha + \beta_1 * (\text{FinDev}_c * \text{FinDep}_i) + \sum \beta_i \text{Industry}_i + \sum \beta_c \text{Country}_c \qquad (1)$$

where Exp_{ic} is the value of the exports of country *c* in industry *i* normalized by country *c*'s GDP (the export share variable); FinDev_c is the level of financial development in country *c*; FinDep_i is the external finance dependence of industry *i*; Industry_i, $i = 1, \ldots, 35$, are industry dummy variables for the 36 ISIC Rev. 2 industries; and Country_c, $c = 1, \ldots, 55$, are the country dummy variables for our fifty-six countries. Industry dummies control for the possibility that some industries might be more likely to engage in exports than other industries for reasons other than the quality of business law. The country dummies control for country-level variables other than the quality of business law, such as GDP, that might affect a country's exports. All the regressions were run with robust standard errors with errors clustered at the country-level. Moreover, the financial development variable was instrumented by a country's legal origins (although this does not qualitatively affect the results).

If financial development improves the export share of externally finance dependent industries *relative* to industries not reliant on external finance, then the coefficient on the interaction of FinDev_c * FinDep_i, β_1, should be positive. In other words, the interaction coefficient should measure the relative impact on externally-finance dependent industries' exports of a change in a country's level of financial development.

The results, consistent with the findings of Beck (2003) and Manova (2006), are reported in Table III. Only the results from the three most commonly-used proxies for a country's level of financial development are presented: Total Capitalization, Private Credit and Market Capitalization. Using all U.S. firms in the 1980s as the basis for computing the external finance dependence of an industry, all three proxies for financial development are statistically significant at the 1% level (Panel A of Table III). The total market capitalization point estimate is 1.22, private credit is 1.25 and market capitalization is .63. Based on these point estimates, the level of private credit in an economy appears to be a more important factor than the size of the equity markets in explaining the export share of a country's externally-finance dependent industries, although both are important. Using young U.S. firms in the 1980s (Panel B of Table III) and, alternatively, all U.S. firms in the 1970s (Panel C of Table III) as the bases for computing industry external finance dependency, one gets similar results: all three coefficients for the three financial development proxies are positive and statistically significant at the 1% level with the private credit point estimate being, as before, substantially larger than that for market capitalization, although with market capitalization still being important.

The variables "Commercial-Central Bank," "Liquid Liabilities," "Turnover Ratio," and "Value Added" were also used as proxies for financial development. The (unreported) results were not qualitatively changed by using these variables as proxies for financial development.

B. Business Law and Comparative Advantage

In this section, we turn to our primary hypothesis. We test whether countries with "higher-quality" business law, as measured by the various law and finance indexes, export more products as a percentage of GDP of externally-finance dependent industries than countries with lower quality business law. Moreover, we also simultaneously test whether countries with "high-quality" business law export relatively fewer products of less externally-finance industries as a result of having "high-quality" business law.

We ran the following regression:

$$\text{Exp}_{ic} = \alpha + \beta_1 * (\text{BusQ}_c * \text{FinDep}_i) + \sum \beta_i \,\text{Industry}_i + \sum \beta_c \,\text{Country}_c \qquad \textbf{(2)}$$

where BusQ_c is the quality of business law in country c. The other variables are the same as in equation (1). Eight proxies where used for the quality of business law: the anti-self-dealing index; ex ante private control of self-dealing index; ex-post private control of self-dealing index; public enforcement index; accounting standards index; anti-directors index; disclosure index; and the creditor rights index.

As before, all the regressions were run with robust standard errors with errors clustered at the country-level. Moreover, to address reverse causation, $BusQ_c$ is instrumented for, as is standard in the law and finance literature, by the legal origins of the country. Therefore, equation (1) is the second stage of a 2SLS estimation. As a robustness check, $TradeSh_{ic}$ was also used as a dependent variable with $TradeSh_{ic}$ being defined as the difference in the exports in industry *i* minus the imports in industry *i* normalized by country *c*'s GDP. The results are not qualitatively affected by substituting $TradeSh_{ic}$ for $Exports_{ic}$ as the dependent variable (unreported regressions). Moreover, all the results reported in Table IV are not affected by whether the legal origins instrument is used or not (uninstrumented results are not reported).

If the quality of business law, as proxied by one of the eight measures, improves the export performance of externally finance dependent industries relative to industries not reliant on external finance, then the coefficient on the interaction of $BusQ_c * FinDep_i$, β_1, should be positive. A positive point estimate on the interaction term indicates that an increase in the quality of a country's business regulation will have a greater positive impact on the export share of externally-finance dependent industries than that of industries less reliant on external finance. Alternatively, the greater the external-finance dependency of an industry the great the impact a change in the quality of a country's business law will be on that industry's export share.

The results, measuring external finance dependency of an industry based on all U.S. firms in the 1980s and using export share as the dependent variable, are reported in Panel A of Table IV. The coefficient estimates on the industry and country dummies are not reported. The explanatory power of the independent variables in equation 2 in accounting for variation in export performance across industries and countries is in the 83% to 85% range. Six of the eight proxies for business law quality are statistically significantly positive with at least 10% statistical significance with the accounting standards, the disclosure index, public enforcement index, and the creditor's rights index all being statistically significant at the 1% level. Of the later four indexes, the accounting standards index and the disclosure index are always statistically significantly positive at the 1% level regardless of the way in which external finance dependency is calculated (Panel B are the results when external dependence is calculated using only young 1980s U.S. firms and Panel C are the results when external dependence is calculated using all U.S. firms from the 1970s). The disclosure index's coefficient point estimate, however, is far larger in all three specifications than that of the accounting standards' estimate.

Mandatory disclosure requirements, as proxied by the accounting standards index and, in particular, the disclosure index, appear therefore to be consistently important, economically and statistically, as a source of comparative advantage in externally-finance dependent industries. There is a variety of mechanisms by which such an association might arise: improved disclosure reducing the adverse selection costs of external finance, reducing private information, and increasing transparency at the firm-level thereby making it more difficult for controlling shareholders and managers to engage in value-destroying or value-transferring activities. Other legal indexes appear to also be important, the public enforcement and the anti-directors index is always significant and positive at the 5% level and the creditor rights and ex post private control of self-dealing index is always significant and positive at the 10% level.

It is worth noting that a benefit of focusing on within country differences – in these results on the difference in the export share of a country's external finance dependent industries and its industries not reliant on external finance – is that it allows one to control for country-level characteristics, sometimes difficult to measure, through the simple device of introducing a country dummy. The country-level characteristics that can be controlled for through a country dummy are those country characteristics that systematically affect export share across all industries, external finance dependent or not. Accordingly, when we included in our regressions indexes such as a "rule of law" index or a "corruption" index as independent variables, the results in Table IV do not change. When the dependent variable is a country-level variable, such as market capitalization, controlling for this kind of country-level characteristic in this way is not possible.

C. Industry Firm Size

Turning to the first question raised in section C of Part III, does an increase in a county's level of financial development disproportionately benefit (in terms of export share) industries where firm size is naturally "small" taking into account the external finance dependence of the industry. Again, such an effect might be the result of small firms having more difficulty raising external finance than large firms in financially undeveloped countries. In order to test this hypothesis, we ran the following regression:

$$Exp_{ic} = \alpha + \beta_1(FinDev_c{*}FirmSize_i{*}FinDep_i) + \beta_2(FinDev_c{*}FirmSize_i) + \beta_3(FirmSize_i{*}FinDep_i) + \beta_4(FinDev_c{*}FinDep_i) + \sum\beta_i\, Industry_i + \sum\beta_c\, Country_c \qquad (3)$$

The coefficient value of interest is β_1, the coefficient on the triple interaction term. If increases in the level of financial development – the $FinDev_c$ variable – enhances the ability of industries with small firm size – as measured by the $FirmSize_i$ variable – to export their goods *relative* to industries with larger firms, taking into account the external finance dependency of the firm's industry, then the point estimate for β_1 should be negative. The expected sign is negative because as the firm size variable increases, firm size in the industry becomes larger.

All the second level interaction permutations were included in the regression so that the triple interaction does not capture their omitted effects. As before, industry and country dummies were included, robust errors clustered at the country level, and all the interaction terms instrumented by legal origins (although instrumenting does not qualitatively change the results). "Total Capitalization" is used as the proxy for financial development as it captures both the size of a country's equity markets as well as the overall private credit provision by financial institutions in the country. We vary the definition of firm size using all five definitions (5, 10, 20, 100, 500) of industry firm size.

Results are reported in Table V. Financial development does appear to disproportionately benefit industries with "small" firm size in terms of their export share using the 5, 10 and 20 definitions of firm size. The triple interaction coefficient point estimate is negative and statistically significant at the 1% for the 5 and 10 definitions of firm size and at the 5% level for the 20 definition of firm size. However, this effect statistically disappears, with the point estimate is still negative but small, when firm size is defined using the 100 or 500 cut-off. Moreover, the size of the effect of financial development on the relative export performance of industries with "small" firm size is monotonically decreasing towards zero as the size cut-off moves from 5 to 500 employees. The progression in the point estimates, moving from the 5 to 500 definition of firm size, is -1.15, -.44, -.18, -.03, and -.01. One possible interpretation of this progression is that industries with a large share of firms above a sufficiently high size threshold do not face different financing constraints than those of industries with a smaller share of firms of the same size threshold. Only industries with a relatively large fraction of truly "small" firms are disproportionately benefited in terms of their export share by increased levels of financial development. This differential effect gradually disappears as one gradually relaxes the strictness of the definition of a "small" firm.

Interestingly, these results parallel findings reported by Beck, Demirgu-Kent, Laeven and Levine (2004). They found that only when one uses a firm size definition based on employment share by firms with 20, 10 or 5 or fewer employees does one get a statistically significant result at the 1% or 5% level on the effect of the provision of private credit in an economy on the average annual grow rate in real value added for "small" firm industries. With a definition of firm size based on the 100 employee cut-off the positive effect of private credit on "small" firm industries has statistical significance at only the 10% level and no statistical significance when firm size is based on the 500 employee cut-off in their study.

What of the hypothesis that a liberalized capital account should enhance the ability of "large" relative to "small" firms to raise external finance taking into account the external finance dependence of the industry? As before, to address this question we again run a regression with a triple interaction:

$$\mathrm{Exp}_{ic} = \alpha + \beta_1(\mathrm{CapAcct}_c{*}\mathrm{Firm\ Size}_i{*}\mathrm{FinDep}_i) + \beta_2(\mathrm{CapAcct}_c{*}\mathrm{Firm\ Size}_i) + \beta_3(\mathrm{Firm\ Size}_i{*}\ \mathrm{FinDep}_i) + \beta_4(\mathrm{CapAcct}_c{*}\mathrm{FinDep}_i) + \sum\beta_i\,\mathrm{Industry}_i + \sum\beta_c\,\mathrm{Country}_c \quad \textbf{(4)}$$

Again, the point estimate of interest is β_1. If increases in capital account liberalization – the $\mathrm{CapAcct}_c$ variable – enhances the ability of industries with large firm size to export their goods *relative* to industries with smaller firms, taking into account the external finance dependency of the industries, then the point estimate for β_1 should be positive. As before, we vary the definition of firm size using all five definitions (5, 10, 20, 100, 500) of industry firm size.[5]

The results are reported in Table VI. Using the strictest definition of firm size, the cut-off of five employees or less, the coefficient value is .92 and statistically significant at the 1% level. The cut-offs of ten and twenty also result in positive and statistically significant at the 1% level coefficients. Firm size using the 100 definition results in a positive point estimate at the 5% statistical significance while the 500 definition is statistically insignificant. Similar to the financial development results, the size of the point estimate is monotonically decreasing towards zero as the size cut-off moves from 5 to 500 employees. The progression in the point estimates, moving from the 5 to 500 definition of firm size, is .92, .43, .2 , .04, and .01.

Finally, turning to the last hypothesis, does higher-quality business law improve the export performance of small, external-finance dependent industries relative to larger, external-finance dependent industries? To test this, we replaced the variable $CapAcct_c$ with $BusQ_c$ in equation 4. As before, we use the eight proxies for the quality of business law.

The results are reported in Table VII. The results do not support the hypothesis. Using the firm size cut-off of ten, six out of the eight proxies for business law quality do not result in statistically significant point estimates, even at the 10% level, for the triple interaction coefficient. The two statistically significant variables, at the 5% level, are when business law quality is proxied for by public enforcement and creditors' rights. Both point estimates are, however, positive rather than negative as the hypothesis would have predicted. No consistent result emerges when one varies the definition of "small" firm.

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[5] Second level interaction terms, industry and country dummies where included and robust standard were clustered at the country level as before.

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TABLE I: CORRELATION BETWEEN MEASURES OF FINANCIAL DEVELOPMENT

	Private Credit	Liquid Liabilities	Commercial Central Bank	Market Capitalization	Value Traded	Turnover Ratio	Total Capitalization
Private Credit	1.00						
Liquid Liabilities	0.81	1.00					
Commercial Central Bank	0.65	0.44	1.00				
Market Capitalization	0.63	0.55	0.56	1.00			
Value Traded	0.68	0.62	0.37	0.62	1.00		
Turnover Ratio	0.55	0.45	0.26	0.38	0.92	1.00	
Total Capitalization	0.95	0.79	0.68	0.84	0.72	0.54	1.00

TABLE II: CORRELATION BETWEEN MEASURES OF QUALITY OF BUSINESS LAW AND CAPITAL ACCOUNT LIBERALIZATION

	Anti-director index	Accounting standards	Anti-self-dealing index	Ex-ante private control of self dealing	Ex-post private control of self dealing	Public enforcement	Disclosure	Creditors' rights	Capital liberalization
Anti-director index	1.00								
Accounting standards	0.51	1.00							
Anti-self-dealing index	0.64	0.45	1.00						
Ex-ante private control of self dealing	0.50	0.29	0.91	1.00					
Ex-post private control of self dealing	0.62	0.52	0.79	0.47	1.00				
Public enforcement	0.23	0.09	-0.08	-0.17	0.07	1.00			
Disclosure	0.58	0.66	0.66	0.47	0.71	-0.10	1.00		
Creditors' rights	0.05	0.17	0.30	0.27	0.24	-0.07	0.08	1.00	
Capital Account Liberalization	-0.28	-0.15	0.01	0.07	-0.08	-0.04	0.00	-0.17	1.00

TABLE III: EXPORT SHARE AND FINANCIAL DEVELOPMENT

Panel A	Dependent Variable: Export Share of an Industry		
(External Dependence All Firms 1980s) × (Fin. Dev.)			
ext. dep. × total capitalization	1.22***		
ext. dep. × private credit		1.25***	
ext. dep. × market capitalization			0.63***
Adj. R-squared	.86	.87	.86
Number of Observations	1492	1982	1492

Panel B	Dependent Variable: Export Share of an Industry		
(External Dependence Young Firms 1980s) × (Fin. Dev.)			
ext. dep. × total capitalization	0.51***		
ext. dep. × private credit		0.44***	
ext. dep. × market capitalization			0.32***
Adj. R-squared	.85	.85	.85
Number of Observations	1409	1874	1409

Panel C	Dependent Variable: Export Share of an Industry		
(External Dependence All Firms 1970s) × (Fin. Dev.)			
ext. dep. × total capitalization	2.97***		
ext. dep. × private credit		2.74***	
ext. dep. × market capitalization			1.75***
Adj. R-squared	.85	.85	.85
Number of Observations	1450	1925	1450

Notes: * = *p*-values < .1; ** = *p*-values < .05; *** = *p*-values < .01

TABLE IV: EXPORT SHARE AND QUALITY OF BUSINESS LAW

Panel A	Dependent Variable: Export Share of an Industry							
Quality of Business Law as Measured by Regulation								
(External Dependence All Firms 1980s) × (Regulation)								
ext. dep. × anti-director index	0.30**							
ext. dep. × accounting standards		0.04***						
ext. dep. × anti-self-dealing index			-0.14					
ext. dep. × ex-ante private control of self dealing				-1.67**				
ext. dep. × ex-post private control of self-dealing					1.14*			
ext. dep. × public enforcement						4.00***		
ext. dep. × disclosure							1.77***	
ext. dep. × creditor's rights								0.56***
Adj. R-squared	.83	,85	.85	.84	.85	.85	.85	.85
Number of Observations	1501	1385	1241	1501	1501	1501	1385	1913

Panel B	Dependent Variable: Export Share of an Industry							
Quality of Business Law as Measured by Regulation								
(External Dependence Young Firms 1980s) × (Regulation)								
ext. dep. × anti-director index	0.20**							
ext. dep. × accounting standards		0.02***						
ext. dep. × anti-self-dealing index			0.46					
ext. dep. × ex-ante private control of self dealing				-0.18				
ext. dep. × ex-post private control of self-dealing					0.81**			
ext. dep. × public enforcement						1.30**		
ext. dep. × disclosure							1.04***	
ext. dep. × creditor's rights								0.23*
Adj. R-squared	.83	.85	.85	.85	.85	.85	.85	.85
Number of Observations	1418	1308	1172	1418	1418	1418	1308	1808

Notes: * = *p*-values < .1; ** = *p*-values < .05; *** = *p*-values < .01

Panel C	Dependent Variable: Export Share of an Industry							
Quality of Business Law as Measured by Regulation								
(External Dependence All Firms 1970s) × (Regulation)								
ext. dep. × anti-director index	1.01***							
ext. dep. × accounting standards		0.13***						
ext. dep. × anti-self-dealing index			1.88					
ext. dep. × ex-ante private control of self dealing				-1.52				
ext. dep. × ex-post private control of self-dealing					4.18***			
ext. dep. × public enforcement						7.75***		
ext. dep. × disclosure							5.27***	
ext. dep. × creditor's rights								1.67***
Adj. R-squared	.83	.85	.84	.84	.85	.84	.85	.85
Number of Observations	1458	1346	1206	1458	1458	1458	1346	1858

Notes: * = *p*-values < .1; ** = *p*-values < .05; *** = *p*-values < .01

TABLE V: EXPORT SHARE, TOTAL CAPITALIZATION, AND FIRM SIZE

	Varying firm size as measured by				
Dependent Variable: Export Share of an Industry	Firm Size 5	Firm Size 10	Firm Size 20	Firm Size 100	Firm Size 500
(External Dependence All Firms 1980s) × (Fin. Dev.) × (Firm Size)					
ext. dep. × total capitalization × firm size	-1.15***				
ext. dep. × total capitalization × firm size		-0.44***			
ext. dep. × total capitalization × firm size			-0.18**		
ext. dep. × total capitalization × firm size				-0.03	
ext. dep. × total capitalization × firm size					-0.01
Adj. R-squared	.84	.84	.84	.84	.84
Number of Observations	1410	1410	1492	1492	1492

Notes: * = *p*-values < .1; ** = *p*-values < .05; *** = *p*-values < .01

TABLE VI: EXPORT SHARE, CAPITAL ACCOUNT LIBERALIZATION AND FIRM SIZE

	Varying firm size as measured by				
Dependent Variable: Export Share of an Industry	Firm Size 5	Firm Size 10	Firm Size 20	Firm Size 100	Firm Size 500
(External Dependence All Firms 1980s) × (Fin. Dev.) × (Firm Size)					
ext. dep. × capital liberalization × firm size	0.92***				
ext. dep. × capital liberalization × firm size		0.43***			
ext. dep. × capital liberalization × firm size			0.20***		
ext. dep. × capital liberalization × firm size				0.04**	
ext. dep. × capital liberalization × firm size					0.01
Adj. R-squared	.84	.84	.84	.84	.84
Number of Observations	1401	1401	1481	1481	1481

Notes: * = *p*-values < .1; ** = *p*-values < .05; *** = *p*-values < .01

TABLE VII: EXPORT SHARE, QUALITY OF BUSINESS LAW AND FIRM SIZE

small firm variable is 10	Dependent Variable: Export Share of an Industry							
Quality of Business Law as Measured by Regulation								
(External Dependence All Firms 1980s) × (Regulation) × (Firm Size)								
ext. dep. × anti-director index × firm size	0.14							
ext. dep. × accounting standards × firm size		0.00						
ext. dep. × anti-self-dealing index × firm size			0.43					
ext. dep. × ex-ante private control of self dealing × firm size				0.46				
ext. dep. × ex-post private control of self-dealing × firm size					0.37			
ext. dep. × public enforcement × firm size						1.71**		
ext. dep. × disclosure × firm size							0.13	
ext. dep. × creditor's rights × firm size								0.25**
Adj. R-squared	.83	.85	.85	.85	.85	.85	.85	.85
Number of Observations	1421	1309	1173	1421	1421	1421	1309	1811

Notes: * = *p*-values < .1; ** = *p*-values < .05; *** = *p*-values < .01

Does FDI cause productivity growth?

The Evidence from China.

By

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Abstract

The role of FDI in generating faster economic growth is a central question in growth theory and continues to be a subject of debates. The endogenous nature of these two variables presents an obstacle for estimating the impact of one on the other. When standard OLS is used to estimate the relationship, it results in biased and inconsistent estimators. However, this endogeneity problem can be overcome by the use of a valid instrument. This paper uses the Chinese government's FDI policy shift of the early 1990s as an instrument for the Chinese FDI in a 2SLS analysis. The results indicate that an increase in the level of FDI as a share of existing capital stock causes an increase in the growth rate of productivity which implies that foreign capital is far more productive than domestic capital after controlling for the province fixed-effects.

Section 1: Introduction

What is the relationship between Foreign Direct Investment (FDI) and productivity growth? The role of FDI in generating faster economic growth is a central question in growth theory and continues to be a subject of debates. The proponents of the "push" theory of FDI argue that FDI inflows to the recipient country promote productivity growth by providing a venue for technology transfer across countries.[1] The supporters of the rival "pull" theory argue that the causal relationship actually goes the other way and claim that, in fact, the high rate of return on investment from existing high productivity growth "pulls in" FDI.[2] Finally, there is a skeptical third group who diligently document the correlation between FDI and productivity growth but are reluctant to make any kind of claims with regard to the direction of the causal relationship between FDI and productivity growth. While agreeing that FDI and productivity are indeed highly correlated, and that it may be that one causes the other, they maintain that the relationship cannot be estimated because FDI and productivity are endogenous.

Existing work in the field provides convincing evidence that that FDI is often associated with higher productivity growth. However, as the skeptics rightly point out, such is merely an evidence of correlation, not causality. Because of lingering endogeneity problems, standard Ordinary Least Squares (OLS) estimations will lead to a biased and inconsistent estimator. Therefore, in order to credibly establish causality, it is necessary to use an appropriate instrumental variable that is highly correlated with one of the endogenous variables but has no contemporaneous relationship to the other. If such a variable were to be found, the IV approach (or 2SLS) could be used to reliably estimate the relationship between the two endogenous variables, FDI and productivity growth.[3]

The purpose of this paper is to propose one such instrument. I estimate the relationship between productivity growth and FDI of China by using the government-dictated change in FDI policy in China of the early 1990s as an instrument for FDI inflows. Not surprisingly, FDI inflows into China and the FDI policy change are highly correlated. Moreover, the policy change affected productivity growth only indirectly, through increased FDI. Therefore, this policy change fits both criteria of an instrumental variable extremely well, and this instrument

[1] See Caves, 1974; Kokko, 1994, Oulton, 1998; Blomstrom and Sjoholm, 1999; Xu, 2000; Liu and Wang, 2003.
[2] See Haddad, 1993; Aitken and Harrison, 1999.
[3] See Kennedy's A Guide to Econometrics, Chapter 8, Fifth Edition, MIT Press, 2003.

allows for an unbiased estimation of the causal relationship between FDI and productivity growth.

China is an obvious choice for a study of the relationship between FDI and productivity.[4] Among developing countries, China is the largest recipient of FDI for the last 30 years, having received over 1.29 trillion U.S. dollars in FDI between 1978 and 2005, an amount that dwarfs any other country's save for the United States[5]. At the same time, the rate of Chinese economic growth during the last three decades has been phenomenal, with an average GDP growth rate of almost 10 percent per annum from 1978 to 2005. This trend is currently expected to continue, at least for the next few years.

In this paper, I aim not only to establish the existence of a causal relationship between FDI and productivity growth, but also to estimate its magnitude by utilizing the previously mentioned instrumental variable. The motivating factor behind this study is the observation that, in China as well as in most other developing nations, technological progress seems to be accompanied by FDI inflows. My main finding is that FDI positively and significantly affects the rate of productivity growth in China even after controlling for the *pulling* effects of productivity on FDI, and that foreign capital is far more productive than domestic capital. While this does not, by itself, invalidate the "pull" theory, I believe it provides strong empirical evidence for the validity of the "push" theory of FDI causing productivity growth.

The paper is organized as follows: section 2 briefly describes related literature, section 3 discusses the data used for analysis, section 4 includes a brief history of FDI in China followed by a discussion of the use of the instrumental variable, and Section 5 describes methodology used for empirical exercise and the results of the analysis. The implications of the empirical findings are finally considered in section 6.

Section 2: Related Literature

The relationship between FDI and productivity growth has been the subject of intense scrutiny by economists in recent years –with China earning a generous amount of attention. Some are mainly descriptive, restricting themselves to simply illustrating the source and distribution of FDI or its correlation to productivity growth.[6] For example, Dayal-Gulati and

[4] Another advantage of studying a single country rather than pursuing a cross-country empirical exercise is that the complications created by cross-country differences do not come into play when using provincial-level data for a single country such as China.
[5] In 2005, GDP of the US was larger than that of EU.
[6] See Kamath (1990), Zhang (2003)

Husain (2002) examine data for Chinese provinces from 1978-1997, and conclude that the inflow of FDI positively correlates with higher levels of per capita income and with increased rates of income growth. They use a variation of Mankiw, Romer, and Weil version of the Solow model to demonstrate that conditional convergence holds for Chinese provinces but the acceleration of economic growth is accompanied by a greater income disparity among provinces. They also find that costal regions attracted more FDI, and they attribute this to the more developed infrastructure in those regions.

Beyond those that simply document a high correlation, there has been some progress in addressing the role and impact of FDI on productivity growth. Starting with Caves (1974), but also Kokko (1994), Oulton (1998), Blomstrom and Sjoholm (1999), and Xu (2000), many studies have found that FDI positively affects productivity growth through promoting technology transfer and diffusion of knowledge across national borders. A host of studies have emphasized how FDI also affects the host country's market structure by facilitating an environment of increased competition resulting in efficiency gains (Liu and Wang, 2002).[7] Grossman and Helpman (1995) cite FDI as an important source of human capital augmentation, technology change, and spillovers of ideas that advance the rate economic growth for the recipient country.

Some studies have attempted to quantify the impact of FDI on productivity growth. Graham and Wada (2001) divide China into 6 regions (North, Northeast, Coastal, Southeast, South, and West) and the data into two time periods (1978-1990 and 1991-1997) and estimate the TFP growth by running an OLS of real GDP growth on capital, labor, and a time trend for each region for each time period. They find that the largest change in TFP growth rate over the two time periods occurred in the Costal region (0.03 to 0.26) followed by Northern (0.00 to 0.18), Northeastern (0.01 to 0.10), Southern (0.12 to 0.14), and Southeastern (0.18 to 0.19). The regression coefficient for the Western region was not statistically significant. They also observe that both Coastal and Northern regions received the most amount of FDI over the 1991-1997 period. With these two findings at hand, Graham and Wada inaccurately conclude that FDI promotes acceleration of TFP growth instead of recognizing that it as a mere incidence of high correlation.[8]

[7] See Hymer (1976), Buckley and Casson (1976), Dunning (1993), and Caves (1996).
[8] Another weakness of Graham and Wada's work is their choice to group 31 provincial level areas into 6 regions. This method disregards cross provincial characteristic differences within these each group.

Other works attempted to measure the impact of FDI on productivity growth by running a simple OLS. Shiu and Heshmati (2006) regress productivity growth on FDI and other variables and conclude that both FDI and information and communication technology investment are significant factors that affect TFP growth. Although they find that a 1% increase of FDI will increase TFP growth by 0.79 percentage point with random affects and heteroskedastic variance assumptions, this result is biased due to the endogenous nature of FDI and productivity growth. In order to combat the inherent endogeneity problem FDI and productivity growth, Chen, Chang, and Zhang (1995) perform a regression of economic growth on lagged FDI and lagged domestic saving. They find that a 1% increase in FDI from a year ago affects the growth of GNP by 0.63 percentage points and conclude that FDI promotes productivity growth. This could be a good workaround for the endogeneity problems, if productivity growth followed a random walk. However, current productivity growth strongly depends on past productivity growth, and therefore past productivity growth and lagged FDI are still endogenous. Therefore, it follows that a simple OLS even with lagged FDI term would still result in biased and inconsistent estimators.[9]

There are also a number of studies on FDI and TFP that are focused on industrial level data. A study by Liu and Wang (2002) investigates the effect of FDI on TFP growth for a cross-sectional sample of Chinese industrial sectors. They regress the log of TFP on log of FDI as a share of capital stock for an industry after controlling for R&D, domestic saving, human capital, and firm size. Their main finding is that a 1% increase in FDI as a share of existing capital would raise productivity by 0.31% or 0.16% if interaction terms are excluded. Zhang and Zhang (2003) and Castejon and Worz (2006) also pursued an analysis of FDI and productivity at the industry level and find support Liu and Wang's results. Unfortunately, these results and conclusions are also biased because FDI and productivity are jointly determined even at the industrial level.

In direct contrast to the works mentioned above, a score of studies suggest that FDI's impact on economic growth may be negative. Chen, Chang, and Zhang (1995) claim that capital inflow is actually harmful to developing countries because FDI tends to substitute domestic

[9] Chen, Chang, and Zhang (1995) include a time trend in their regression. They suggest that this time trend inclusion "avoid(s) a potential spurious relation caused by a common trend observed in GNP, domestic savings and FDI". However, this still does not solve the endogenous issue between GNP and FDI.

saving instead of supplementing it. Moreover, several researchers also contend that a huge inflow of FDI can exacerbate the hosting country's balance of payments deficits as debt repayment obligations rise in proportion to the amount of investment received. For example, studies by Haddad (1993), Aitken and Harrison (1999), and others document an inverse relationship between FDI inflow and productivity growth.

Section 3: Data

Organization of Chinese provinces/municipalities:

China is divided into 31 regions. There are 27 provinces and 4 municipalities in direct control of the Chinese central government. Chongqing became a separate municipality in 1997. Prior to that time, it was a part of Sichuan province. Therefore, Sichuan data is inclusive of Chongqing's share prior to its separation. Although data for Sichuan and Chongqing are separately available from 1997, Sichuan and Chongqing are treated as one region for the purposes of this paper for data comparability.

<u>31 Regions are:</u>

1) Beijing	11) Zhejiang	21) Hainan
2) Tianjin	12) Anhui	22/23) Chongqing/Sichuan
3) Hebei	13) Fujian	24) Guizhou
4) Shanxi	14) Jiangxi	25) Yunnan
5) Inner Mongolia	15) Shandong	26) Tibet
6) Liaoling	16) Henan	27) Shaanxi
7) Jilin	17) Hubei	28) Gansu
8) Heilongjiang	18) Hunan	29) Qinghai
9) Shanghai	19) Guangdong	30) Ningxia
10) Jiangsu	20) Guangxi	31) Xinjiang

Data Requirements:

For the analysis of this paper, provincial level productivity and FDI data is used. In order to estimate productivity growth, I use data on GDP, capital stock, labor, capital share of income, and labor share of income. My model also relies on a few essential assumptions, discussed below. Since capital stock data is not available, it must be estimated from gross capital formation data with a particular depreciation assumption and an initial stock of capital. Provincial level CPI is necessary for deflating nominal values. The nominal exchange rate is used to convert the nominal FDI data that is reported by the Central government in US dollars.

GDP, Labor and other variables:

GDP, Labor, and other control variables are compiled from various years of *China Statistical Yearbook* and provincial yearbooks. Chinese government reports two measures of nominal GDP: GDP by expenditure approach and GDP by income approach. Both are available at the provincial/municipal level. Interestingly, no statistical or measurement errors exist between these two variables.[10] Data is available from 1952 to present for all but four provinces. Jiangxi, Guangdong, Hainan, and Sichuan data is available from 1978. Nominal GDP data was deflated using Chinese CPI with 1978 as the base year because GDP deflator data at the provincial level is not available. Labor data is available from 1978 to present for all regions from the same source.[11] Labor is defined by the Chinese central government as the total number of employed (adults) persons in units of ten thousand.

CPI:

Regional CPI data is available from different years from provincial yearbooks. 22 out of 31 Chinese regions report CPI data from 1952 to 2004. Of the 9 remaining areas, Tibet and Ningxia are excluded from the analysis due to lack of all necessary data, including the CPI data. Of the 7 regions with missing CPI data, Anhui, Jiangxi, Henan, and Gansu are missing a few data points (see below), from 1952 to 1971. However, the remaining three, Guangxi, Heinan, and Sichuan, are missing 20+ years of CPI data. For all 7 regions, all missing CPI data occur prior to 1978. Whenever available, regional CPI is used to convert nominal data, but for those regions with missing CPI data, national CPI is used as a substitute in interpolating missing data. Since regional CPI does not differ significantly from one region to another prior to 1978, national CPI is an acceptable proxy.

<u>7 Regions with missing CPI data are</u>:

1) Anhui: Missing 1952 to 1959 & 1967 to 1970
2) Jiangxi: Missing 1968 to 1970
3) Henan: Missing 1952 to 1956
4) Guangxi: Missing 1952 to 1977 (with a few sporadic entries)
5) Heinan: Missing 1952 to 1978
6) Sichuan: Missing 1952 to 1977 (with a few sporadic entries)
7) Gansu: Missing 1966 to 1971

[10] See Young (2003) footnote 5 for further discussion on data manipulation.

[11] Data prior to 1978 are mostly not available. If available, then it is usually recorded every 5 years.

Foreign Direct Investment Data:

FDI data is obtained from *Comprehensive Statistical Data and Materials on 50 Years of New China* published by the Chinese Central Government in 2006. Most of the data is now available electronically via China Data Online, a paid subscription service provider.[12] This is the only known vendor authorized to sell electronic Chinese government data by the National Bureau of Statistics of the Chinese central government. Although comprehensive in Chinese, the availability of monthly and yearly data at the provincial level in English is limited.

Each province/municipality annually reports the total amount of foreign capital for that year to the central government. Along with the aggregate data, three sub-categories (Foreign Loans, Foreign Direct Investment, and Other Foreign Investments) are reported as well. FDI data is consistently available only for 29 out of 31 areas of China for the 1980-2004 period.[13] FDI data is published in the units of millions of U.S. dollars by the Chinese government. Therefore, the exchange rate published by the IMF is used to convert the data into Chinese yuan. Afterwards, the nominal values of FDI are deflated using Chinese CPI for the analysis.[14]

China Statistical Yearbook provides more detailed data of FDI than *Comprehensive Statistical Data and Materials on 50 Years of New China*. It further subdivides FDI data into 6 categories: 1) joint venture enterprises, 2) cooperative operation enterprises, 3) foreign investment enterprises, 4) foreign investment share enterprises, 5) cooperative development, and 6) others.[15] However, none of the six categories of data are available consistently for all regions for all years. Therefore, aggregate FDI data was chosen for the analysis.

Capital Stock:

The national capital stock as well as one for each region must be estimated from the available capital formation data because neither the regional nor the central governments of China published it. Physical capital stock is a function of existing capital stock at the beginning

[12] Prior to September 2006, only printed version of provincial level data were available. As for some of the older data, they are scanned and save into files, so they can't be read by other programs. Although the raw data is in electronically format, they are not truly "electronic".

[13] Chongqing became an independent municipality in 1997. Prior to 1997, it was a part of Sichuan province. Tibet and Ningxia are excluded.
Therefore, Sichuan's FDI data is inclusive of Chongqing's share up to 1996.

[14] The base year for CPI is 1978, the most commonly used base year in related literature. Whenever provincial CPI is not available, national CPI was used as a proxy.

[15] For exact definition of each category in English, see Fung et al (2002), pages 8-10.

of the period (K_t), the depreciation rate (δ), and the amount of new investment added over the period (I_t).

$$K_{t+1} = f(K_t, \delta, I_t) \qquad \textbf{(1)}$$

If a linear relationship is assumed, then

$$K_{t+1} = (1-\delta)\, K_t + I_t \qquad \textbf{(2)}$$

Solving backwards, equation (2) becomes,

$$K_{t+1} = (1-\delta)^t K_0 + (1-\delta)^{t-1} I_0 + (1-\delta)^{t-2} I_2 + \ldots + (1-\delta)\, I_{t-1} + I_t \qquad \textbf{(3)}$$

In order to accurately estimate the capital stock of China, three components are necessary: the stream of capital formation/new investment (I_t), depreciation rate (δ), and initial capital stock (K_0). Obtaining a reliable initial capital stock of China is not possible, and therefore, an exact calculation of capital stock of China is not achievable. However, as t increases, the first term on the RHS of the equation (3) becomes negligible because $(1-\delta)^t$ asymptotically approaches zero. For example, if δ is equal to 4% and t is equal to 55 years, then $(1-\delta)^t$ is about 5%. On top of a small discounting factor, $(1-\delta)^t$, if the level of initial capital stock also happens to be small—which is quite possible given China's GDP level in 1955 relative to that of the present time— then, assuming the first term to be zero is not detrimental to capital stock estimation. In other words, if enough years of data are available—*t* is very large—then a good approximation of capital stock is possible without knowing the exact level of initial capital stock.

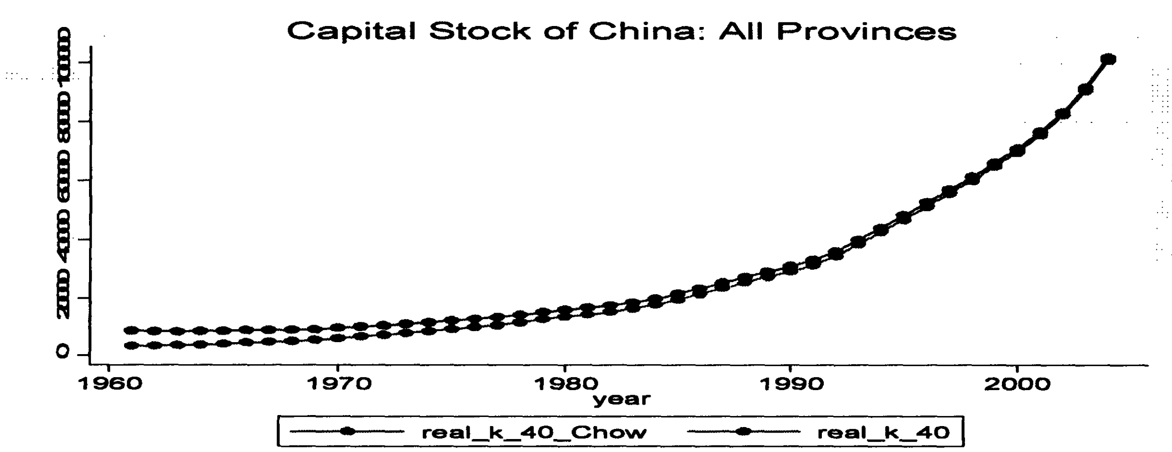
However, the assumption of zero initial capital stock can lead to a greater than actual capital stock growth rate for the beginning years of the sample period starting in 1952. To combat this issue, Young (1995), Hall and Jones (1999), and Unel and Zebregs (2006) calculate the initial level of capital stock by $K_0 = I_0 / (g + \delta)$ where g is the annual growth rate of capital stock and δ is the depreciation rate assumed to be 5%. I adopt the same methodology for the initial level of capital for all provinces. However, given I only use the 1980 – 2004 period in this paper and that FDI data is only available from the early 1980s, this technical issue has an insignificant overall effect on the results of this paper.

Various years of *China Statistics Yearbook* and *Provincial Yearbooks* report Gross Capital Formation data as a part of Gross Regional Product. The provincial capital formation data goes back to 1952, so almost 55 years of data are available for the estimation. Data is available from 1952 to 2005 for all but 4 regions.[16] Jiangxi, Guangdong, Hainan, and Sichuan are missing capital formation data from 1952 to 1977. Regional CPIs were used to convert nominal values to real ones.[17]

For robustness checks, I compare my capital stock estimation with that of Chow (1993). Chow's seminal work uses publicly available official data from various years of *China Statistics Yearbook* along with confidential government data to estimate the capital stock of China for 1952 to 1998. Obviously, the main advantage of using his estimation for comparison comes from the extra accuracy derived from his use of classified government data. He was able to obtain national level land and inventory data, going back to the early 1960s and incorporated them into his capital stock estimation. The availability of confidential data may allow for a more accurate calculation of capital stock. However, Chow's methodology is not directly applicable because first, the classified data is not in the public domain, and second, even if it was, it is not available at the provincial level according to Chow. Still, Chow's data can serve as an empirical robustness check to validate the theoretical assumptions of this paper.

First, I compare Chow's capital stock estimation with mine at the national level.[18] Both measures of capital stock for the country are illustrated in Figure 1. Although Chow's estimation for Chinese capital stock is not identical to my calculation, both estimates are quite close overall, and the difference is not statistically significant. Prior to early 1980s, Chows estimate of capital

Figure 1: Capital Stock Estimation (units in 100,000 1978 yuan)



[16] Ningxia and Tibet are also missing I data. However, these two provinces are completely excluded from all analysis due to lack of all available data, so missing I is irrelevant here.
[17] See CPI description section for further information.
[18] Available at http://www.princeton.edu/~gchow/downloads.html. Click on "Data Files on Chinese Economy" link.

stock is slightly larger than mine. For example, in 1978, Chow's estimate of Chinese national capital stock is 1.411 billion in comparison to my calculation of 1.145 billion of 1978 yuan. However, the difference between two calculations of capital stock prior to 1980's is unimportant since FDI data is available from early 1980s for most of the regions at the earliest. By early 1980s –which is where my data analysis starts—the difference between the two is negligible, and by 2000, it is all but identical.[19]

In order to compare my provincial level capital stock estimation with that of Chow, I calculate a second capital stock variable that utilizes Chow's national capital stock estimation because Chow's estimation is not available at the regional level.[20] Chow's 1978 capital stock value is distributed among 31 regions by using each region's GDP as a share of national GDP in 1978. Using 1978 data point as the anchor for each region, I derive a new capital stock variable for each region with regional capital formation data and CPI starting in 1952.[21] The comparison of my capital stock estimation and the one based on Chow's 1978 value as an anchor is illustrated in Appendix 1 for each of 31 regions. Even at the provincial level, the result is similar to that of the national level. For all regions, the difference between the two capital stock estimation is insignificant for the relevant period from the early 1980s and almost identical by 2000.

Productivity Growth:

Total output is a function of capital, labor, and productivity. Data for output, capital, and labor is readily available, but productivity is not because it can not be easily measured. Since three out of four variables are known, if a specific functional form is assumed, then, it is possible to estimate or "back out" the productivity variable by utilizing the other three known variables. Let Y be output, X be capital stock, L be labor, β be capital's share and $(1-\beta)$ be labor's share of income. Let the production function take the following form:

$$Y = A_t \, X^{\beta} L^{(\theta)} \tag{5}$$

where X is defined as the sum of domestic capital stock (K) and foreign capital stock (K*), and α is a measure of relative productivity of foreign capital to domestic capital.

$$X = K + \alpha \, K^{*} \tag{6}$$

[19] Chow's estimate is available only through 1998.
[20] As for the depreciation rate, Chow (1993) uses 4% and many of the literature (see China Report: Social and Economic Development 1949-1989, China Statistical Information and consultancy Service Center, 1990) uses 4.5%. This seemed a bit low, so I calculated the capital stock variable using a range of 4% to 8% but the change in the series is statistically insignificant.
[21] Base year is 1978.

If X is observed then, the Solow Residual also known as the TFP growth is defined as

$$S = \frac{\Delta Y}{Y} - \beta\left(\frac{\Delta X}{X}\right) - \theta\left(\frac{\Delta L}{L}\right) \tag{7}$$

However, $\overset{\frown}{X} = \mathrm{K} + \mathrm{K}^*$ is observed instead of X, leading to

$$\overset{\frown}{S} = \frac{\Delta Y}{Y} - \beta\left(\frac{\Delta \overset{\frown}{X}}{\overset{\frown}{X}}\right) - \theta\left(\frac{\Delta L}{L}\right) \tag{8}$$

From equation (6), it can be derived that

$$\frac{\Delta X}{X} = \frac{\Delta K}{X} + \alpha\left(\frac{\Delta K^*}{X}\right) \quad \text{and} \quad \frac{\Delta \overset{\frown}{X}}{\overset{\frown}{X}} = \frac{\Delta K}{\overset{\frown}{X}} + \left(\frac{\Delta K^*}{\overset{\frown}{X}}\right) \qquad \textbf{(9), (10)}$$

If $\frac{K}{X}$ is close to $\frac{K}{\overset{\frown}{X}}$ and $\frac{K^*}{X}$ close to $\frac{K^*}{\overset{\frown}{X}}$ then, substituting (9) and (10) into (8) establishes a relationship between measured TFP growth and FDI.[22]

$$\overset{\frown}{S} = A + (\alpha - 1)\left(\frac{\Delta K^*}{\overset{\frown}{X}}\right) = A + (\alpha - 1)\left(\frac{FDI}{\overset{\frown}{X}} - \frac{\delta K^*}{\overset{\frown}{X}}\right) \tag{11}$$

where $\Delta K^* = FDI - \delta K^*$. If foreign capital is more productive than domestic capital, then α is great than 1, and equation (11) implies that the higher the ratio of FDI to measured capital stock net of depreciation of foreign capital stock, the higher the measured productivity growth. This is the equation that will be estimated using the 2SLS methodology with a new instrument.

Section 4: Policy Change as an Instrument for FDI

Brief History of FDI in China

The Communist Chinese central government, established in 1949, naturally distrusted foreign capitalists. The prevalent sentiment within the Chinese Communist Party at that time was that foreign investors had come to exploit and sully China by hiring cheap labor, extracting natural resources on the cheap, and gravely polluting the environment (Chow 2007). This anti-FDI sentiment reached its apex during the Cultural Revolution started in 1966 by Chairman Mao.

[22] The proportion of FDI to capital stock is relatively small for China allowing for this assumption.

By the end of the Cultural Revolution in 1976[23], China was almost completely closed to all foreign investments. Soon after the death of Chairman Mao, Deng Xiaoping became the leader of the Communist Party in 1978. Under Deng's leadership, China adopted significant changes to its FDI policies. As "the chief architect of China's [economic and political] reforms"[24], Deng championed opening up China's economy to the rest of the world.(Zhao, 1993). In 1979, as a direct result of Deng's leadership, the Fifth National People's Congress passed regulations allowing legal status for FDI for the first time in China's post-war history.

Deng's efforts of aggressively attracting FDI and his open-door policy ("kaifang zhenze" in Chinese) continued throughout the 1980s. In 1980, four Special Economic Zones (SEZs) were established in Shenzhen, Zhuhai, Shantou, and Xiamen. These zones were provided with infrastructure necessary for attracting FDI. The power supply was guaranteed, at very low cost, even when the rest of the country experienced devastating power shortages. The communication and transportation infrastructure required by foreign investors was put in place and paid for in full by the central and provincial governments. In addition, the central government allowed for special business laws and favorable tax conditions for foreign investors in these zones (Chow 2007). Moreover, the Sixth National People's Congress in 1982 formally adopted the decision to open China to the rest of the world, even modifying the Chinese constitution in the process. In 1983, the Party proceeded to further liberalize the Chinese domestic market. In 1984, twelve coastal cities were designated as Technology Promotion Zones, and in 1985, three "development triangles" were established near the Yangtze River delta, the Pearl River delta, and the Min Nan region in Fujian. In 1986, wholly foreign-owned enterprises were allowed, and Hainan Island became the fifth SEZ in 1988.(Chow, 2007) In 1990, the Pudong District of Shanghai became a special development zone.[25]

Although many regulations and policies favorable to FDI were put in place during the 1980s, it was Deng's famous Southern Tour (Nan Xun Jiang Hua) of 1992 that "reignited the radical economic reforms"(Shambaugh, 1993). As "the most powerful man in China,"[26] Deng

[23] Chairman of the Communist Party of China, Mao Zedong, started the Cultural revolution on May 16, 1966. Although Mao officially declared the Revolution to be over in 1969, some Chinese historians assert that it was not over until the arrest of the Gang of Four, led by Mrs. Mao, in 1976.
[24] Renmin Ribao, April 3, 1992 (Chinese).
[25] See appendix for documentation in Chinese.
[26] Deng officially stepped down as the leader of China in 1989, but he still retained most of his political power throughout the 1990s.

toured several southern cities and SEZs, concluding his tour in Shanghai (Zhao, 1993). During his trip, Deng highlighted the accomplishments of economic reform (results of the open-door-policy) of the 1980s but re-emphasized the need for further reform.[27] Immediately afterwards, he championed his economic vision during the Fourteenth People's Party Congress in October 1992 (Whiting, 1995). As a direct result of Deng's southern tour and his later efforts, in a massive shift in policy, the central government of China made the paradigm-changing decision to allow the inflow of FDI to be regulated by provincial governments and other forms of local government in 1992.[28] Although many political obstacles for FDI had been eliminated prior to 1992, due to the central government's inefficient bureaucratic nature it was not until this shift of decision making power from the central to the provincial governments in 1992 that attracting large amounts of FDI became a real possibility for China. This shift of power—the single most important change in policy— made China a very attractive destination for FDI by eliminating the need to deal with the often unfathomable and insurmountable layer of Chinese central government bureaucracy.

Policy change of 1992-1994 as an instrument for FDI:

As could be expected of such a large country, the shift of power to the regional/provincial governments was not uniform across all regions.[29] Some provinces and municipalities experienced a greater freedom to accept FDI than others. Out of 27 provinces and 4 municipalities in direct control of the Chinese central government, 9 provinces and 3 municipalities (from this point on these will be called *"affected"* regions) benefited far more than the rest from the events of 1992-1994. While elsewhere the progress in openness was haphazard at best, in these areas 15 free trade zones, 32 economic and technological development zones, and 53 industrial development zones were established during that period as a result of the policy change. The 12 regions that benefited the most significantly from the policy change are: Beijing, Tianjin, Hebei, Liaoling, Shanghai, Jiangsu, Zejiang, Fujian, Shandong, Guangdong, Guangxi, and Hainan.

[27] Deng Xiaoping was quoted by Beijing Review as saying, "Although some people are opposed to the way the SEZs are run, nobody can deny the great, proven achievements of the SEZs." Beijing Review, 35;15 (April 13-19, 1992), pg 4.

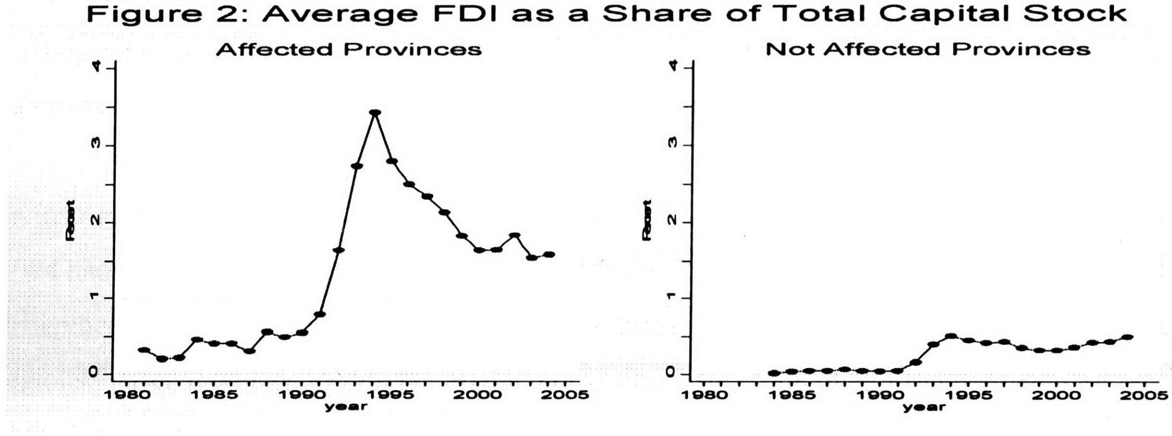
[28] Jin, Qian, and Weingast (1999) show that the new fiscal contracting system, that allowed local governments to engage in long-term investment contracts and keep 100% of marginal revenue, is associated with "faster development of non-state enterprises and more reform in state enterprises." Dayal-Gulati and Husain (2003).

[29] See appendix 2 for partial documentation in Chinese. Full documentation available upon request.

The policy change of 1992-1994 is an ideal instrument choice for FDI because the FDI policy change is not contemporaneously correlated with productivity growth. It is very important to note that the Chinese government did not target the 12 fastest growing regions for the FDI policy change and implementation. The twelve affected regions are in fact an odd mix of some already developed and some underdeveloped provinces. Since the Chinese government does not openly discuss policy issues, the reasons why the government targeted the particular 12 regions it did are still being debated in some circles. However, why some were chosen is not so enigmatic. Both Beijing and Shanghai are politically important in China, and both were already among the most developed areas in China at the time. Targeting Guangdong is also in line with the expectation of massive FDI flows into the region after nearby Hong Kong returns to China in 1997. But others—such as Guangxi, and Liaoling— were at the time relatively underdeveloped so they were clearly not chosen based on their past performance, so the policy "treatment" can safely be assumed to have little to do with past productivity growth.

Its high correlation with FDI inflow is another reason for choosing the policy change as an instrument for FDI. Unlike the government's efforts of the past, the policy change of 1992-1994 had a massive impact on FDI inflows into China. Figure 2 below illustrates the significant difference in the amount of real FDI as a share of capital stock between affected and unaffected regions.[30] Between 1992-1994, the affected provinces as a group experienced an increase of about 2 percentage points on average in comparison with 0.3 percentage point average increase for the unaffected group. In contrast, the FDI-friendly polices put in place between 1980-1992 had minimal, if any, impact on FDI for either group.

Figure 2: Average FDI as a Share of Total Capital Stock



[30] See appendix 2 and 3 for FDI/K by province.

Difference in Difference Analysis:

In order to quantify the difference between "affected" and "unaffected" regions, I calculate the impact of the policy change on FDI. Table 1 illustrates the Difference-in-Difference estimation of the effect of the central government's FDI policy change on the ratio of FDI to existing capital stock for all provinces. The mean of FDI share for each group for each time period are appropriately listed in corresponding cells, and standard errors are listed in parenthesis below each mean. FDI as a share of existing capital is the main variable of interest as was discussed previously in equation (11).

The Difference-in-Difference estimator was calculated over 4 different time periods. The top panel of Table 1 displays the estimation result from limiting the estimation range to 5 years prior to and after the policy implementation of 1992. Before the policy change, the level of FDI as a share of existing capital for the affected provinces was 0.35 percentage points higher than that of unaffected provinces on average. However, the difference widens to 1.66 percentage points after the policy change. For the affected provinces, there was an increase of 1.6 percentage point over 10 years, but the change for the unaffected provinces was only 0.29 percentage point. Thus, there was a 1.31 percentage point increase in the difference of the ratio of FDI to capital stock between affected and unaffected states; this is the difference-in-difference estimate of the policy's impact on the level of FDI as a share of capital stock.

Table 1 also lists Difference-in-Difference estimates for longer time periods. When the estimation range is widened to 7, 9, and 12 years on either side of the policy change date of 1992, estimates decrease to 1.25, 1.11, and 0.98 percentage points respectively. All estimates are significant at the 1% level, and thus it attests to the quality of the "policy" variable as a good instrument. This result is consistent with Figure 2 of the previous section. For the first few years, as a direct result of the policy change, the affected provinces experienced a dramatic increase in FDI as a share of measured capital stock. However, within 8 years or so, FDI shares for most of the affected provinces stabilized at a level that is higher than pre-1992 levels, but much lower than the peak experienced immediately after the policy change. In contrast, unaffected provinces as group did not experience the same fluctuation. Instead, the slight increase in FDI as a share of capital stock immediately following the policy change maintained its level throughout the post-policy period for the unaffected provinces. Therefore, the Difference-in-Difference estimator is smaller as the estimation range increases.

Table 1. Difference-in-Difference Estimates of the Impact of Policy on FDI	Before Policy	After Policy	**Time Difference**
A. The policy impact on FDI/K (1992 +/- 5 years)			
Affected States	0.392	1.991	1.599
	(0.658)	(0.226)	(0.236)
Not Affected States	0.047	0.336	0.289
	(0.012)	(0.034)	(0.037)
Location Difference:	0.345	1.655	**1.310**
	(0.090)	(0.294)	**(0.239)**
B. The policy impact on FDI/K (1992 +/- 7 years)			
Affected States	0.363	1.911	1.548
	(0.054)	(0.176)	(0.186)
Not Affected States	0.043	0.344	0.301
	(0.009)	(0.030)	(0.033)
Location Difference:	0.319	1.567	**1.247**
	(0.074)	(0.226)	**(0.187)**
C. The policy impact on FDI/K (1992 +/- 9 years)			
Affected States	0.329	1.734	1.405
	(0.045)	(0.133)	(0.147)
Not Affected States	0.039	0.331	0.292
	(0.008)	(0.024)	(0.030)
Location Difference:	0.290	1.403	**1.113**
	(0.065)	(0.172)	**(0.143)**
D. The policy impact on FDI/K (1992 +/- 12 years)			
Affected States	0.328	1.629	1.301
	(0.045)	(0.108)	(0.134)
Not Affected States	0.039	0.365	0.326
	(0.008)	(0.027)	(0.038)
Location Difference:	0.289	1.264	**0.975**
	(0.065)	(0.140)	**(0.120)**

Note: Each cell contains the mean for the group identified. Standard errors are listed below in parenthesis. Differences-in-Differences results are in bold.

Section 5: Two-Stage Least Squares Analysis

The econometric model to be estimated in this paper for region i at time t is:

$$(\text{TFP Growth})_{it} = \beta_i + \theta \left(\frac{FDI_{it}}{\hat{X}_{it}} - \frac{\delta K_{it}^*}{\hat{X}_{it}} \right) + \varepsilon_{it} \qquad \textbf{(12)}$$

and θ is the estimator of interest in the structural equation because it captures the impact of FDI as a share of measured capital stock on productivity growth as was derived in equation (11). However, as discussed before, it cannot be estimated directly by a simple OLS of equation (12) because productivity growth and FDI are endogenous. Therefore, a two-stage least squares method with an instrument will be used instead.

First Stage: FDI/K on Policy Change

First Stage: $$\left(\frac{FDI_{it}}{\hat{X}_{it}} - \frac{\delta K_{it}*}{\hat{X}_{it}}\right) = \lambda_1 + \lambda_2(\text{Policy})_i + \gamma_i\ (\text{province dummies})_i + \eta_{it} \qquad \textbf{(13)}$$

The first stage of 2SLS estimation consists of regressing the ratio of FDI to existing capital stock net of foreign capital depreciation, one of the endogenous variables, on the instrument called "policy" along with province dummies to obtain the fitted value to be used as an explanatory variable in the second stage (or the reduced form) regression of equation (12). The binary "policy" variable was constructed by assigning a value of 1 to those provinces affected by the policy change, and a value of 0 to those that were unaffected[31]. Although the "policy" variable works well as an instrument for net FDI/K, the LHS variable of equation (13), it does not capture all information embedded in the FDI/K variable because all of the numerical information from FDI data gets compressed into an indicator variable that takes a value of either 0 or 1.

Table 2 below summarizes the first-stage regression results with standard error listed in the parentheses below the estimated coefficient for different estimation periods.[32] For the estimation period over the 5 years immediately before and after 1992, the coefficient on the policy variable is 0.58 and is statistically significant at the 1% level. This suggests that, on average, the net ratio of FDI to capital stock was about 58 basis points higher for the affected provinces after controlling for state fixed effects. When the estimation range is increased to 7, 9, and 12 years, the coefficient decreases to 0.39, 0.29, and 0.23 percentage points respectively.

Table 2. 2SLS First Stage

	policy +/- 5 yrs (1988-1997)	policy +/- 7 yrs (1986-1999)	policy +/- 9 yrs (1984-2001)	policy +/- 12 yrs (1981-2004)
Coefficient	0.577	0.385	0.285	0.226
S.E.	(0.192)	(0.134)	(0.097)	(0.080)

[31] As discussed in Section 4, "unaffected" does not mean that a particular region was insulated from the FDI policy change. Rather, "unaffected" regions were not specifically targeted by the Central government for dramatic policy change, and therefore, were affected much less by it.

[32] All variables are measured in real terms with the base year of 1978.

This result is in line with the findings of the difference-in-difference estimation results of the previous section. Because the first stage regression uses the change in foreign capital stock as defined by $\Delta K^* = FDI - \delta K^*$ instead of FDI as a share of measured capital, the magnitude of the coefficients are not directly comparable to that of the Difference-in-Difference estimation coefficients. However, in both cases, as the estimation range increases, the impact of the policy on FDI as a share of measured capital stock becomes less pronounce for the affected provinces, and therefore, the coefficients are progressively smaller.

Second Stage: FDI and Productivity

Second Stage: $(\text{TFP Growth})_{it} = \beta_i + \theta \left(\frac{FDI_{it}}{\hat{X}_{it}} - \frac{\delta K_{it}^*}{\hat{X}_{it}} \right) + \gamma_i (\text{province dummies})_i + \varepsilon_{it}$ **(14)**

The fitted values obtained from the first-stage regression were used as the explanatory variable in the second-stage regressions. Table 3 below summarizes the results from regressing TFP growth on the ratio of net FDI to measured capital stock for different estimation periods. For the estimation period over the first 5 years immediately before and after 1992, the coefficient θ on the explanatory variable is 0.88 and is statistically significant at the 1% level. This suggests that a 1 percentage point increase in net FDI as a share of capital stock increases the growth of TFP by 0.88 percentage points after accounting for province fixed effects. As the estimation range varies between 5 to 12 years from the policy change date of 1992, the value of coefficient θ varies between 0.68 and 0.94 percentage points, and all are significant at 1% level.

Table 3. 2SLS Second Stage

	policy +/- 5 yrs (1988-1998)	policy +/- 7 yrs (1986-1999)	policy +/- 9 yrs (1984-2001)	policy +/- 12 yrs (1981-2004)
Coefficient	0.875	0.678	0.781	0.936
S.E.	(0.456)	(0.367)	(0.358)	(0.346)

The second stage regression results suggest that foreign capital is more productive than domestic capital. From equation (6) recall that total capital stock was defined as $X = K + \alpha\ K^*$. The coefficient α captures the relative productivity differential between domestic and foreign capital stock. α greater than 1 would suggest that foreign capital is more productive than domestic capital stock. Since θ of equation (14) is equal to (α -1) of equation (12), a positive θ would imply that α is greater than 1. Regardless of the estimation range, θ is estimated to be always greater than 0 in the second stage regression. Therefore, α is indeed greater than 1 ranging between 1.68 to 1.94 percentage points, and it can be concluded that foreign capital is at least 1.68 times more productive than domestic capital.

Section 6: Conclusion

The main contribution of this paper to the existing literature is the introduction of a variable that can successfully instrument for FDI, one of the endogenous variables. The instrumental variable is derived from the Chinese foreign policy shift of the early 1990s that affected FDI inflows to different regions of China differently, with implementation probably based on political considerations but not on previous productivity growth in those provinces. To this author's knowledge, the use of the FDI policy shift of the early 1990s as an instrument for FDI in establishing a relationship between FDI and productivity has not been pursued previously. This variable is an excellent instrument for FDI because it is not contemporaneously correlated with productivity but highly correlated with FDI. The results from my difference-in-difference analysis support the use of the policy change as an acceptable instrument. Unlike the policies of the past, the FDI policy change had a large impact on FDI inflows into those provinces that were affected by the policy change. The 2SLS analysis strongly supports the "push" theory claim that an increase in FDI as a share of capital stock causes a significant impact on the growth of productivity after controlling for province fixed effects, and that foreign capital, namely FDI, is at least 1.68 times productive than domestic capital.

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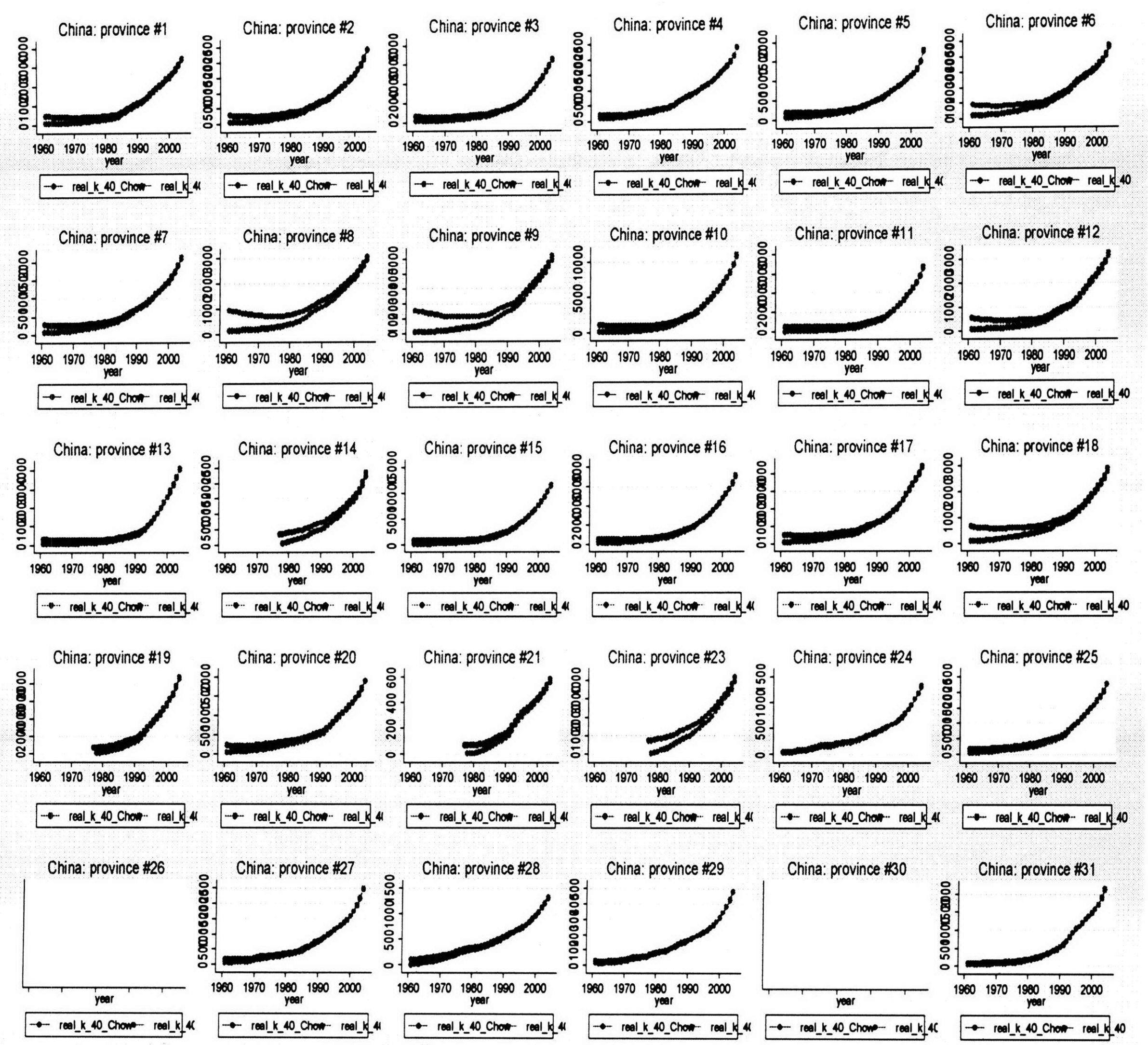
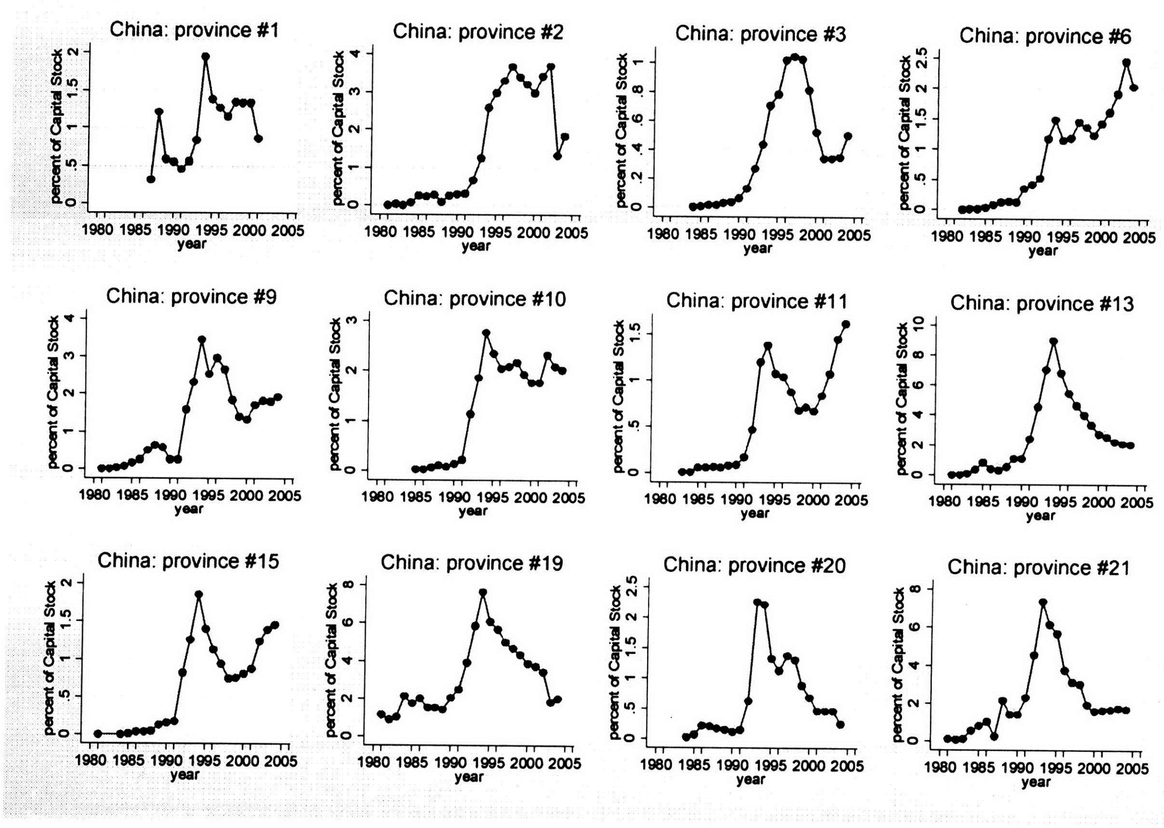
Appendix 1: Capital Stock Estimation Comparison for 30 regions.



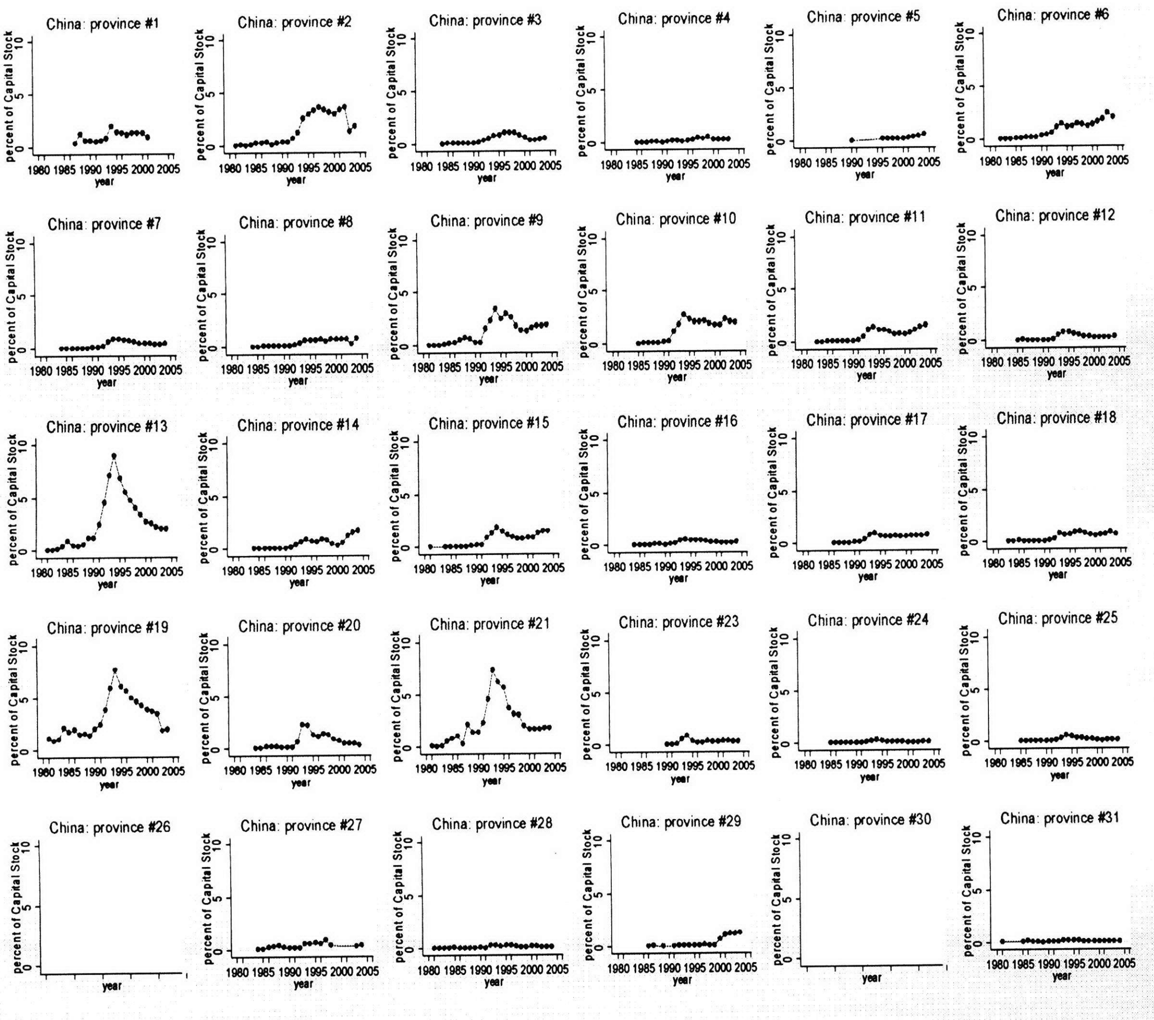
Appendix 2: FDI as a Share of Capital Stock of 12 Affected Regions[33]



Appendix 2 (above) captures the magnitude of difference between "affected" and "unaffected" regions. A graph of FDI as a share of GDP for each of 30 regions (drawn to scale) is shown below as Appendix 3.[34] For all 30 regions, FDI as a share of capital stock increased at or around 1992. However, the increase for "affected" regions (provinces 1, 2, 3, 6, 9, 10, 11, 13, 15, 19 ,20, 21) is much larger in comparison to that of "unaffected" regions. Therefore, it is clear that the policy aim of 1992 was successful in attracting FDI into selected regions.

[33] Province numbers correspond to the list on section 3 of this paper.
[34] As noted earlier in the paper, Chongqing and Sichuan are treated as one province.

Appendix 3: Real FDI as a Share of Capital Stock[35]



(As mentioned in the text, both Tibet and Ningxia (province #26 and #30) do not have enough data, so were excluded from all analysis.)

[35] Province numbers correspond to the list on section 3 of this paper.