HOW MUCH DO LOW WAGES MATTER FOR FOREIGN INVESTMENT?

The Case of China*

Xuepeng Liu
Department of Economics
Syracuse University
xuliu@maxwell.syr.edu

Mary E. Lovely
Department of Economics
Syracuse University
melovely@maxwell.syr.edu

Jan Ondrich
Center for Policy Research
Department of Economics
Syracuse University
jondrich@maxwell.syr.edu

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Abstract

Although studies of aggregate investment flows provide consistent evidence that capital is attracted to low wages, there is almost no empirical support for this proposition from plant-level location choice studies. We examine the provincial location choices of firms investing in China during 1993-1996 to offer two main contributions. First, using data on 2884 manufacturing equity joint venture (EJV) projects, we investigate the extent to which standard estimation suffers from omitted variable bias. Applying a two-step technique, developed by Petrin and Train (2004), to conditional logit analysis, we find strong support for the attractiveness of low wages. Our estimates indicate a downward bias of 50-120 percent in the wage coefficients estimated with standard techniques. Second, we find that low wage locations are more attractive to unskilled-labor-intensive plants than to skill-intensive plants, although this effect is significant only for investors from OECD countries. The attraction of low wages for investors from ethnically-Chinese-economies, in contrast, is sensitive to the intensity of competition from other low-income countries for exports to the United States. This study provides the first estimates of how skill intensity and competition for export markets influence the probability a multinational firm will choose a given location.

I. Introduction

The entry of China, India, and the former Soviet Union into the world economy has, in effect, doubled the global labor force. China alone accounts for more than half of this increase. In most developed countries, average real wages have lagged well behind productivity gains. A common explanation for lagging wages is that this massive entry of cheap labor into the international trading system has reduced the bargaining power of workers.[1] Faced with globally mobile capital, workers are forced to compete on the basis of wages or else watch while capital and jobs move offshore. These pressures are reported to be particularly pronounced in labor-intensive activities and in sectors with low entry barriers.

While much attention is focused on the impact of wage competition on developed country workers, the gradual dismantling of barriers to trade in labor-intensive sectors, such as the Multi-Fiber Arrangement, has left few Western workers engaged in these activities. Competition for investment capital in these industries is largely occurring among the developing countries. For example, in apparel imports to the United States, China's most important competitor is Mexico, with many subcontractors active in both countries.[2] Competition to provide cheap labor is thought to extend to regions within the two countries, as each local labor market seeks to attract foreign capital through low wages.[3] For China, Anita Chan (2003) notes that Guangzhou and Shenzhen, the first cities opened to foreign investment and those with the highest income, maintain a ratio of minimum wage to average income that falls below the central

[1] The Economist, "China and the world economy," July 30, 2005, p. 62.
[2] Each country supplies around 15 percent of the American import market.
[3] Robert J.S. Ross and Anita Chan (2002, p. 9) forcefully articulate this view. Jagdish Bhagwati (2004, p. 129) provides a response.

government guideline of 40 percent. In her view, these low minimum wages are a response to the threat to coastal competitiveness posed by lower wages in the Chinese interior.

While this type of anecdotal information suggests a struggle for mobile capital, our understanding of the extent to which foreign capital is attracted to low-wage locations is surprisingly incomplete. Although studies of aggregate investment flows provide consistent evidence that capital is attracted to low wages, there is almost no empirical support for this proposition from studies that use microdata. Such data are prized because aggregate data often is not rich enough to explore key questions such as how factor intensity influences wage sensitivity. As a result, the empirical record is virtually silent on some of the most important controversies surrounding international capital flows.

This study offers two main contributions: first, it suggests an econometric explanation for the failure of previous microdata studies to find a deterrent effect of high wages on FDI and illustrates the use of a method to correct it; second, it provides new evidence on the nature of firm's attraction to low wages. Using data on the location choices of 2884 manufacturing equity joint venture (EJV) projects in China during 1993-1996, we investigate the extent to which standard estimation techniques suffer from omitted variable bias. We apply a two-step technique, developed by Petrin and Train (2004) to estimate unbiased price elasticities for differentiated products, to the standard location-choice analysis. Using this control function approach, we estimate a downward bias of 50-120 percent in the wage coefficients estimated with standard techniques. Second, we find that low wage locations are more attractive for unskilled-labor-intensive activities than for skill-intensive activities, although this effect is significant only for investors from OECD countries. The attraction of low wages for investors from ethnically-Chinese-economies, in contrast, is sensitive to the intensity of competition from other low-income countries for exports to the United States. This study provides the first estimates of how skill intensity and competition for export markets influence the probability a multinational firm will choose a given location.

II. Wages, Firm Location Choice, and Omitted Variable Bias

As the literature on FDI flows is large, our review of previous studies of wages and firm location choice is necessarily targeted.[4] From all studies using aggregate FDI flows, we report only results obtained in studies of investment into Chinese provinces. Among project-level studies, we examine results using data from foreign investment into the United States, the European Union, and China. These regions receive the largest shares of foreign investment and permit study of location choice in the context of centralized labor market regulation.

Table 1 provides a summary of recent studies of the distribution of aggregate FDI flows among Chinese provinces or regions. In all studies except one, the wage is found to be a statistically significant, negative determinant of the value of FDI. This result appears to be robust to the choice of method and to the inclusion of controls for skill level or skill availability. The only study that did not find a significant coefficient for wage, Gao (2002), divides the value of investment among 14 source countries. Despite this exception, these aggregate studies strongly support the view that firms seek locations with low wages, *ceteris paribus.*

Given the uniformity of results from aggregate flows, it is surprising that studies using project-level data do not typically find a significant deterrent effect of wages. An insignificant wage coefficient has been estimated in studies using data on firms investing in the United States (e.g., Ondrich and Wasylenko (1993), Head, Ries, and Swenson (1999), List and Co (2000), and

[4] We also note that there are studies that use cross-country variation in wages, such as Wheeler and Mody (1992) and Wei (2000). Such studies do not find evidence of a race to the bottom in wages. We do not attempt to compare their results with the within China studies, but it is an interesting topic for future research to unify the evidence from both cross-country and within-country studies.

Keller and Levinson (2002)); in Europe (Devereux and Griffith (1998), Head and Mayer (2004)); and in China (Head and Ries (1993)). Indeed, as shown in Table 2, in some specifications the estimated wage coefficient is positive. A recent exception to this pattern is Amiti and Javorcik (2005), who relate changes in the number of firms to changes in the average wage. Given the strong theoretical presumption that wages should be negatively related to location attractiveness, the inability to estimate a significant, negative coefficient is frustrating to many researchers. Head and Mayer (2004) express this frustration and its frequency when they note, "The absence of a significant negative effect of regional wages on location choice in all specifications is disappointing. However, the result is not out of line with other studies…"

When wages and unobserved location characteristics are not independent, standard econometric techniques that require exogenous covariates produce biased estimates. As proposed by Berry (1994) to explain low price elasticity estimates in differentiated product studies, sellers will typically receive higher prices when their product has more desirable omitted characteristics. These omitted characteristics may include any attribute that affects the true value of the product to the buyer. When independence is maintained, buyers look less price-sensitive than they are because they receive more for the price they pay than the econometrician takes into account.[5] Applying this logic to the FDI context, omitted location characteristics that influence worker productivity and wages could lead to biased estimates of the wage sensitivity of investors. If the unobserved factors are otherwise mean independent of observed factors, there is unambiguously a downward bias in standard estimates – firms look less sensitive to the wage than they really are.

[5] Petrin and Train (2004) provide many examples from studies of differentiated product models, including the well-known study by Berry, Levinsohn, and Pakes (1995).

The need to control for unobserved location-specific attributes is widely recognized in studies using repeated cross-sections. Most researchers include geographic fixed effects to address the problem (*e.g.*, Head and Mayer (2004), Keller and Levinson (2002)).[6] However, the use of standard fixed-effect techniques may not be sufficient. Assuming the unobserved attributes are time invariant, there may be insufficient variation over time or too many empty cells to use fixed effects defined over the same geographic unit as the choice set. Keller and Levinson (2002), in their study of FDI inflows to U.S. states, Head and Mayer (2002), in their study of flows to regions within European countries, and Head and Ries (1996), in their study of inflows to Chinese provinces, use regional fixed effects rather than state, country, or provincial fixed effects for this reason. A second reason why fixed effects may not sufficient is that the unobservable attributes may not be time invariant.

Another approach to spatially correlated errors is to estimate a nested logit model. A key assumption underlying the conditional logit model estimates, the independence of irrelevant alternatives, has been shown by Dean, Lovely, and Wang (2005) not to fit the Chinese case well. The nested logit procedure allows location choice to be a two-level nested decision--choosing among Chinese regions and then making a specific choice of province with a region.[7]

As demanding of the data as these procedures are, neither approach fully accounts for the omission of location characteristics correlated with the wage. Indeed, all of the logit studies summarized in Table 2 include either regional fixed effects or use a nested logit procedure. As an alternative, Berry, Levinsohn, and Pakes (1995) develop an approach that is similar in spirit to the inclusion of fixed effects for each choice unit, but which recognizes the need for parsimony.

[6] As Head, Ries, and Swenson (1999) note, this provides a convenient way to capture common attributes. Many studies observe fewer than 1,000 investments and as they sometimes span a decade or more, there are few observations in many year-location cells. Consequently, parsimony is necessary given data limitations.

[7] Further discussion of the application of these methods to modeling firm location decisions can be found in Ondrich and Wasylenko (1993).

This approach, known as the product-market control approach, has been widely used in estimating differentiated product models. It involves estimation of a set of controls that match observed to predicted markets shares. Petrin and Train (2004) identify a number of advantages of this approach, but note that sampling error in market shares enters the estimation equations in a non-linear manner. Unless sampling error in the market shares is minimal, this estimator is not consistent and asymptotically normal. Because the sampling error is unknown for the data we employ in the present study, we choose not to use the product-market control method.

An alternative two-stage method is proposed by Petrin and Train (2004), based on control functions. A control function is a term or factor added to the estimation to control for endogeneity. The control-function approach is a two-step estimation. In the first step, IV regression is used to estimate the variables that enter the control function. This first step basically requires the construction of an expected wage for each province in each year, conditional on all exogenous factors observed by the econometrician. In the second step, the likelihood function is maximized with the control function added in the form of additional explanatory variables. In comparisons to price elasticities estimated without correction and with correction using the Berry, Levinsohn, and Pakes (BLP) method, Petrin and Train find that estimated elasticities are very similar across the control function and product-market control approaches, but that they both differ significantly from the uncorrected estimates. Because of the ease of application and concern about sampling error in market shares, we use the control function approach in this study to correct for omitted variable bias.

III. A Location Choice Model

We use a familiar model in which heterogeneous investors choose a location from among 28 Chinese provinces in which to invest. The model specifies the probability that a province yields the highest profits for a particular investor. As in prior work, we choose functional forms that lead to a final specification that is log linear in the parameters.

The Profit Function

A multinational firm seeks to invest one unit of capital in the form of an equity joint venture (EJV) somewhere in China.[8] The firm will locate in the province that maximizes its profit. The firm produces a good with a generalized Cobb-Douglas technology, using variable inputs of labor, imported inputs, and a vector of intermediate (locally-provided) services. Log profits for firm *i* in province *j* can be written as:

$$\ln \pi_{ij} = \ln(1-\tau_j) + \ln(p_{ij} - c_{ij}) + \ln D_{ij}(p_{ij}), \tag{1}$$

where τ_j reflects the (perhaps concessionary) tax rate on foreign investment in province *j*, p_{ij} is the price charged by firm *i* in province *j*, c_{ij} is the unit cost of producing in province *j*, and $D_{ij}(p_{ij})$ is demand, conditioned on price. Because China is used as an export platform, the prices of many of the good produced by multinational enterprises are given by world markets and demand for the product does not vary by province. This type of FDI is largely associated with investors from ethnically-Chinese economies (hereafter called “Chinese” investors). Investment by Japanese or Western partners (hereafter called “Foreign” investors) is characterized as locating to serve local markets. In this case, p_{ij} depends on local conditions. An alternative to the assumption of a fixed world price is to follow Head, Ries, and Swenson (1999) and assume that demand depends on price, local income, *I*, and a demand shock:

$$\ln D_{ij} = \eta_I \ln I_j - \eta_p \ln p_{ij} + e_{ij}^d. \tag{2}$$

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[8]We take the decision to produce abroad, as well as the region in which the project will be located, as made in a prior stage. Zhang and Markusen (1999) consider the firm’s choice of producing at home and exporting or producing abroad. We also use a static model of the investment decision, as is common in the literature.

In this framework, firms manufacture a unique variety of a differentiated product and they select a price that is a markup over average cost, c: $p_{ij} = (\eta_p /(\eta_p - 1))c_{ij}$, where for profit maximization, $\eta_p > 1$. Combining (1) and (2) and using this pricing rule, the profit function is

$$\ln \pi_{ij} = \ln(1-\tau_j) + \eta_I \ln I_j - \nu_p \ln c_{ij} + \nu_p \ln \nu_p - \eta_p \ln \eta_p + e_{ij}^d, \tag{3}$$

where $\nu_p = \eta_p - 1$. Cost is a function of the wage, *w*, the price of imported inputs, *v*, a price index for locally-provided inputs, $\tilde{p}_s$, and an idiosyncratic cost shock:

$$\ln c_{ij} = \theta_L \ln w_j + \theta_v \ln v_j + \theta_s \ln \tilde{p}_{sj} + e_{ij}^c, \tag{4}$$

where θ_i denotes a cost share. Using (3) and (4), we obtain a standard profit function

$$\ln \pi_{ij} = \alpha + \ln(1-\tau_j) + \eta_I \ln I_j - \nu_p \theta_L \ln w_j - \nu_p \theta_v \ln v_j - \nu_p \theta_s \ln \tilde{p}_{sj} - \nu_p e_{ij}^c + e_{ij}^d. \tag{5}$$

Here, we have grouped together all terms that do not vary by province.

Equation (5) leads to a proposition that motivates the use of interaction terms in our empirical specification.

PROPOSITION 1.

The effect on profits of an increase in the wage:

(i) is larger for firms with a larger labor cost share;

(ii) is larger for firms facing higher demand elasticity.

PROOF: The proportionate change in profit for a proportionate change in the wage is given by:

$$\frac{\hat{\pi}_j}{\hat{w}_j} = -\nu_p \theta_L < 0. \tag{6}$$

The proposition follows directly from the fact that this coefficient is larger in absolute value for larger values of θ_L, labor's cost share, and for larger values of η_p, the price elasticity.

Agglomeration and Local Suppliers

Previous research has shown that foreign firms have a strong tendency to locate in areas where other foreign firms have located. We incorporate agglomeration into our model by adapting the Head and Ries (1996) framework for the localization economies. Head and Ries argue that agglomeration in China is the result of localization economies from concentrations of intermediates providers. They assume the market for local services is monopolistically competitive and that foreign firms use a composite of these services. They show how the equilibrium number of intermediate suppliers depends on the final-good price, the number of foreign firms to which they may sell, N_j^f, and the number of domestic firms who may undertake the costly upgrading necessary to serve foreign firms, $\bar{N}_j^s$. Local service provision requires local labor as an input. Dean, Lovely, and Wang (2005) use this framework to derive an intermediates price index. This index, $\tilde{p}_s$, appears in the profit function (5) and measures the price per effective service unit. Assuming log-linear functional forms, this index can be expressed as[9]

$$\ln \tilde{p}_{sj} = \ln A + \mu_L \ln w_j + \mu_p \ln p_j + \mu_f \ln N_j^f + \mu_s \ln \bar{N}_j^s, \quad (7)$$

where A is a constant and the coefficients are functions of the underlying final-goods and intermediates production parameters. Substituting this expression back into the firm's profit function (5) yields an expression that can be used as the basis for estimation.

Benchmark Estimating Strategy

Our basic estimating strategy is similar to procedures used in most of the studies summarized in Table 2. We use these results as a benchmark for comparison to results obtained using the control function method.

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[9] See Dean, Lovely and Wang (2005) for the derivation based on the Head and Ries (1996) model.

The profit function (5) and the price index (7) yield a linear function for log profits with arguments given by the vector

$$\mathrm{X} = [\ln w, \ln v, \ln(1-\tau), \ln N^f, \ln \bar{N}^s, \ln I]. \tag{8}$$

Letting the error vector $e = e^d - \nu_p e^c$, we obtain $\Pi = \mathrm{X}\beta + e,$ where β is the vector of parameters to be estimated. Our estimation strategy depends on the distribution of the unobserved idiosyncratic terms, e_{ij}. If these features are distributed independently according to a Type I Extreme Value distribution, then the probability, P_k, that province *k* is chosen where *k* is a member of choice set *J* is given by

$$P_k = \frac{\exp\left(x_k \beta\right)}{\sum_{j\in J} \exp(x_j \beta)}. \tag{9}$$

The probability in equation (9) is a conditional logit. This model is well suited to the location choice framework since it exploits extensive information on alternatives, can account for match-specific details, and allows for multiple alternatives.[10] Regional fixed effects are added to the list of regressors to capture regional correlation in the supply and demand shocks.

A Control Function Approach

Despite the inclusion of regional fixed effects, possible endogeneity of the wage remains and can be illustrated by specifying the error in the profit function as a two-component error:[11]

$$\varepsilon_{ij} = \xi_j + e_{ij}. \tag{10}$$

ξ_j is location specific, observed by workers and firms but not by the researcher. e_{ij} is a firm-specific idiosyncratic error, that is assumed to be independent across firms and locations.

[10] An alternative approach is to use count data and a Poisson or negative binomial specification. These count approaches are appropriate when there is a preponderance of zeros and small values for counts (Greene, 2003). Data used by Keller and Levinson (2002) have this characteristic.

[11] This discussion adapts the discussion of consumers' choice among differentiated products in Petrin and Train (2004) to the location choice context.

Defining $\mathbf{X}_j$ as in (8) and letting $\mathbf{Z}_j$ be a vector of instrumental variables, under certain regularity conditions the wage can be expressed as an implicit function of all factors taken as given at the time of the decision:

$$w_j = w_j(\mathrm{X}_j, Z_j, \xi_j). \tag{11}$$

Because wages will be higher in locations with more desirable omitted characteristics, ε_{ij} and w_j will be correlated even after conditioning on $\mathbf{X}_j$.

As proposed by Petrin and Train (2004), the control function approach includes a proxy in the estimating equation to test for and correct the omitted variables problem. In the location choice context, we add to the firm's profit function terms that proxy for the unobserved location characteristics. The method proceeds in two steps. The first step is a linear regression of wages (w_j) on exogenous variables (X_j) and instrumental variables (Z_j) using provincial level data. We use this regression to construct the expected wage for each province in each year, conditional on all exogenous factors observed by the econometrician, and the residual. We use this residual to form the control function, $f(\mu_j, \lambda)$, where μ_j is the residual from the first-stage regression and λ is a vector of estimated parameters. The profit function for firm *i* locating in province *j* can now be written as: $\ln \pi_{ij} = \alpha + x_{ij}\beta + f(\mu_j, \lambda) + (\xi_j - f(\mu_j, \lambda)) + e_{ij}$. The new error, $\eta_{ij} = \xi_j - f(\mu_j, \lambda) + e_{ij}$, includes the difference between the actual province-specific error ξ_j and the control function used in the estimation. Because the control function includes predicted values, the standard errors are incorrect. As described in the Appendix, we use bootstrapping methods to correct the reported errors.

This approach requires an instrument for the first-stage wage regression that is correlated with the wage paid by EJVs, but uncorrelated with firms' location choices, conditional on other exogenous variables. Identifying a suitable instrument requires characterization of the wage setting process in China. As discussed in Chan (2003), while local governments set minimum wages, private firms are otherwise free to set wage levels. Given this, we consider the wages paid by foreign firms as determined by the supply of and demand for labor within the province. Our first stage regression, therefore, is a reduced-form wage equation with controls for labor supply (*e.g*. population, share of labor force with secondary education or more) and for labor demand (*e.g*. the rate at which output of state-owned enterprises (SOEs) is falling, cumulative foreign investment, and the number of local enterprises).

The instrument we use in the first stage is the average industrial wage paid by state-owned enterprises. The SOE wage and the wage paid by EJVs are correlated because all enterprises must consider the local living standard in setting wages. Working in a SOE is an alternative to working in a private firm. Thus, the SOE wage acts as a reservation wage when workers negotiate with private firms. As shown in Appendix Table 1, the first stage regression explains 86 percent of the variation in the private wage (our measure of the average wage paid by EJVs) and the SOE wage is highly significant. The SOE wage is a strong instrument in that adding it to the first stage explains an additional 5 percent of the variation.

We rely on the nature of the SOE wage setting process and SOE productivity-wage gaps to argue for the independence of these wages from unobserved factors that drive foreign firm productivity and, hence, location choice. In China, SOE wages prior to 1996 were largely determined by the central government, despite several sets of wage reforms. Starting in 1985, the Ministry of Labor (MOL) provided some profit-oriented incentives to SOEs, but to a very limited extent. For example, as shown in Yueh [2004], the State Council in 1992 permitted SOEs to set their internal wage structure within the confines of a wage budget established by the government. However, the central government still retained controlled wages through two channels. If a wage bill exceeded the MOL standard, the enterprise paid a Wage Adjustment Tax of 33 percent. Alternatively, the enterprise could propose a wage budget and then submit it for approval to the MOL and the Ministry of Finance (MOF). Deeper reforms of China's SOE wage structure were not implemented until the Ninth Five Year Plan (1996-2000). Therefore, during the time frame of our sample, SOE wages were largely set by central government guidelines.

Evidence from SOE productivity-wage gaps also supports the view that SOE wages do not reflect local attributes that influence foreign-firm productivity. According to official statistics, the average SOE wage has risen faster than inflation since 1980 and faster than nominal productivity. The wages of private and foreign firms have risen at an even faster rate since 1985, but these wage gains have been more than offset by gains in productivity.[12] There are many reasons to explain this gap, including concerns about income inequality (Gordon and Li, 1999), and the possibility of insider control (Aoki, 1995 and Qian, 1995).

Once first-stage residuals are obtained, they are used to form the control function. In several applications of their methods, Petrin and Train (2005) specify the control function as a polynomial of μ_j. We use includes the residual and its square, although we typically find that only the first-order term is significant. This specification is motivated by the observation that the wage in private enterprises reflects the marginal value product of labor. Therefore, workers in provinces with positive residuals have higher productivity than expected, given the exogenous factors observed by the econometrician. Standard theory of the firm suggests that firms choose location based on productivity-adjusted wages. Because the residual provides an estimate of the

[12] Parker (1995) finds that, "In 1992, state industrial wages were 43 percent higher than those available in urban collectives, and only 22 percent below those of the other ownership forms; these workers in other ownership forms, however, were 130 percent (in 1990 prices) to 200 percent (in 1980 prices) more productive than those under state-ownership."

value of labor in that location, conditional on all observed factors, we treat the residual as a productivity measure and enter it into the conditional logit in the same way as the wage.

IV. Data Description and Sources

The sample of equity joint venture investments was compiled by Dean, Lovely, and Wang (2005). The sample contains EJVs undertaken during 1993-1996 using project descriptions available from the Chinese Ministry of Foreign Trade and Economic Cooperation (MOFTEC).[13] Provinces are grouped into five regions: coastal, northeast, central, southwest, and northwest.[14] Figures 1 and 2 provide the distribution of the EJV sample across provinces by source and by skill intensity, respectively. Figure 1 shows that both Chinese and Foreign partners engage in equity joint ventures in all provinces. Investment into the southern coastal region is predominantly Chinese, reflecting the geographic proximity and early opening of these provinces. Investment into the northern coastal region is split more equally between both sources. In Figure 2, we show the distribution of EJVs across province by skill ratio. Most provinces receive investment by firms in all skill-intensity groups. The most prominent specialization occurs in the northwest region, where natural-resource based activities dominate. Shanxi and Ningxia have predominantly low and medium skilled EJVs, and Qinghai Province has only low skilled EJVs.

Our theoretical framework implies the use of the covariate vector **X** given by (8). A complete description of all variable definitions and sources is provided in Table 3. The *Chinese Statistical Yearbook* (various years) was used to compile data on labor supplies, agglomeration,

[13] Equity joint ventures are limited liability companies incorporated in China, in which foreign and Mainland Chinese investors hold equity. For further details, see Fung (1997). Wang (2001) provides additional details on the legal framework for foreign investment.

[14] Coastal: Beijing, Fujian, Guangdong, Hainan, Hebei, Jiangsu, Shandong, Shanghai, Tianjin, Zhejiang; Northeast: Heilongjiang, Jilin, Liaoning; Central: Anhui, Henan, Hubei, Hunan, Jiangxi, Shanxi; Northwest: Gansu, Inner Mongolia, Ningxia, Qinghai, Shaanxi, Tibet, Xinjiang; Southwest: Guangxi, Guizhou, Sichuan, Yunnan.

intermediates suppliers, infrastructure and incentives. Summary data for provincial characteristics are provided in Table 4.

The wage measure (w) is the average provincial wage paid by private and foreign enterprises, drawn from Branstetter and Feenstra (2002). This wage is paid by privately controlled firms. These data are the best available measures of the average wage paid by equity-joint venture firms in each province and year. We also draw from this source the average wage paid by state-owned enterprises, which we use as a first-stage instrument. Wage measures are deflated by a national price deflator to create an average real provincial wage. Average wages do not control for provincial variation in labor quality, so we also include the share of the provincial labor force that has completed senior secondary school or above.

We do not have direct measures of the cost of imported inputs (v) nor the corporate tax rate (τ). To control for provincial variation in these factors, we include an incentive dummy that takes a value of one if there is a special economic zone (SEZ) or open coastal city (OCC) in the province. This variable does not vary during the 1993-1996 period. We also include measures of provincial infrastructure, which influence the local cost of imported inputs. Transport infrastructure is proxied by the length of roads adjusted for provincial size. Telecommunications infrastructure is proxied by the number of urban telephone subscribers relative to population.

The number of foreign firms (N^f) is measured as the real value of cumulative FDI, which we refer to as agglomeration, for the period 1983 to the year before the project is undertaken.[15] Availability of potential suppliers of intermediate goods $(\bar{N}_S)$ is measured by the number of local firms. This measure was created by Dean, Lovely, and Wang (2004) and it was created by taking the total number of enterprises at the township level and above (thereby

[15] Calculated using data from Coughlin et al., 2000.

capturing larger enterprises that may have the capacity to supply a foreign-invested plant) and subtracting the number of enterprises that are wholly or partly foreign owned.[16]

To control for market demand, we include the population of the province and several measures of provincial income. The income measure is the size of the provincial private market, calculated as the private share of output multiplied by provincial GDP. We use non-state output to gauge the size of the market open to foreign enterprises because domestic sales in a province will be limited if demand is substantially satisfied by the state sector. Additionally, to allow for a flexible form for this market measure, we also include the square of this variable. Sales may also be affected by the extent to which a province is liberalizing, so we also include the change in state ownership, measured as the difference in the share of industrial output produced by SOEs between time *t* and time *t*-1.

V. Results

Benchmark and Control Function Results

Table 5 reports the conditional logit results for the full sample. All variables are lagged one year to represent predetermined information, available to an investor at the time of the location decision. The first two panels report results estimated using standard conditional logit analysis. All covariates have the expected signs and all except road density are highly significant. The overall fit of the equation is good and it is comparable to other studies using similar procedures.

For both specifications, we estimate a negative and highly significant coefficient for private wages. In comparing results in the first and second panels, we see that the estimated value of the wage coefficient drops by 44 percent when regional fixed effects are included. Coefficients for other covariates also change in value, especially the coefficients on agglomeration, number of

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[16] Specifically, those firms classified as "foreign-funded" or "funded by entrepreneurs from Hong Kong, Macao, or Taiwan" were subtracted from the total.

local firms, population, and private market size. The regional coefficients indicate that EJVs are more likely to locate in any region other than the Southwest, although the difference is not significant for the Northwest region. As expected, these coefficients are largest for the Central and Coastal zones, which have received the largest share of foreign investment.

The third and fourth panels provide results estimated using the control function approach as well as regional fixed effects. The reported standard errors were corrected using bootstrapping techniques described in the appendix. When the residual from the first-stage wage regression is added to the conditional logit, its coefficient is highly significant and positive, as expected. The wage coefficient remains negative and highly significant. However, it increases substantially in absolute value, indicative of a downward bias in the standard method. The coefficient of -2.12 estimated with the control function is 120 percent larger in absolute value than the coefficient of -0.96 estimated without the control function. Adding higher orders of the residual to the logit changes the quantitative results only marginally, as illustrated by the results in the fourth panel.

Table 6 provides estimated own and cross wage elasticities, by province. The elasticities of province *j* were calculated by $e_j^{own} = \beta^w(1-P_j)$ and $e_j^{cross} = -\beta^w P_j$ where β^w and P_j are the estimated wage coefficient and predicted probability that an investor chooses province *j*.[17] In the left panel, we show elasticities calculated without correcting for omitted variable bias. The right panel shows the elasticities calculated using the logit estimated using a control function. The corrected own-wage elasticities tend to be twice as large as the uncorrected elasticities. These results provide strong support for the view that wage elasticities estimated using the standard conditional logit approach are subject to downward bias from omitted variables.

[17] The cross elasticity of province j is the same for any province *k* due to the IIA assumption of conditional logit.

Looking across locations, the own-wage elasticity is smallest for those provinces with the highest predicted probability of being chosen, including Beijing, Guangdong, and Jiangsu. These provinces, conversely, have the largest predicted cross-wage effects, implying that a decrease in their wage has a larger affect on other provinces than the effect other provinces have on them. These estimates imply a dynamic that differs somewhat from the view expressed by Chan (2003), who fears that coastal provinces maintain low wages to fend off competition from interior provinces. Our estimates indicate that the coastal provinces are less likely to lose investment to other provinces when their wages rise, but have the largest effect on other province's chances of attracting investment if they do attempt to keep wages low.

Allowing for Differential Response by Source and by Skill Intensity

Wage pressures are widely believed to be most severe in those industries that make intensive use of unskilled workers. Indeed, according to Proposition 1, we expect higher wages to have a larger effect on profits in labor-intensive industries and, thus, we expect these industries to be more responsive to provincial variations in labor costs when choosing a location for a joint venture. To test this and to estimate the associated probability elasticities, we allow firm wage sensitivity to vary by skill intensity.

To characterize industry factor intensity, we use information from the United States. Industrial skill intensity is calculated using SIC87 4-digit industry data from the NBER-CES Manufacturing Industry Database. Skill intensity is defined as the number of non-production workers divided by the number of production workers in each ISIC 3-digit industry. The mean skill intensity for the projects in our sample is 0.45. The skill intensity for each industry is provided in Table 9. The most skill-intensive industry is the manufacture of professional, scientific, and controlling equipment, which has a skill ratio of 0.98, and the least skill intensive is textiles, which has a skill ratio of 0.17. Given this measure, we predict that the estimated coefficient for the interaction of the private wage and skill intensity will be positive: wage sensitivity should be lower for more skill-intensive ventures.

We find strong evidence that the attraction of low wages is a function of firm factor intensity. As shown in the first panel of Table 7, the interaction of the log wage and the skill intensity is positive and highly significant, as predicted, for the full sample. Again, the estimated wage coefficient is much larger when the control function is added to the conditional logit estimation. The wage coefficient for an EJV with the mean skill intensity is -1.76. The range of response is wide, however, with an estimated coefficient of -2.07 for the textile industry, -1.60 for electric machinery, and -1.17 for beverages.

Detailed descriptions of Chinese inward investment describe investment from the ethnically Chinese economies of Hong Kong, Singapore, and Taiwan as locating in China to use it as a low-wage export platform. In contrast, investment from Japan, Europe, and the United States is characterized as locating in China to serve local markets.[18] This distinction is viewed as consistent with the greater clustering of overseas-Chinese funded ventures, which are less evenly distributed across provinces than is investment from Japan and the United States.[19] Moreover, evidence provided by Huang (2003) indicates that although Chinese and Foreign investments are similarly distributed across industries, foreign-funded firms make more intensive use of engineers, managers, and college graduates than do overseas Chinese funded firms.[20] This suggests that even within 4-digit industries, there are factor intensity differences associated with source country. For these reasons, we investigate the extent to which these two types of investors differ in their response to wage levels, including an interaction between wage and

[18] This distinction is drawn by Henley, Kirkpatrick, and Wilde (1999).
[19] Huang (2003) provides the standard deviation values for project number and value, by source (p40, n67).
[20] See especially Huang (2003), Table 3.3, p. 134.

factor intensity.[21]

As shown in the second and third panels of Table 7, the probability of a Chinese or a Foreign investor locating in a given province is negatively affected by the provincial wage and this response is highly significant for both groups. The responsiveness of both groups is affected by industrial factor intensity; the interaction terms for both groups are positive and statistically significant. There are differences in the two groups, however, and these differences suggest important differences in the behavior of the two groups. The wage coefficient for Foreign EJVs is -3.81, compared to -2.15 for Chinese investors. However, the effect of factor intensity is twice as large for Foreign investors, 1.67 versus 0.84. Despite this positive interaction, Foreign investors are more responsive than Chinese investors to wages regardless of skill intensity. At the mean skill intensity, the wage coefficient for Foreign investors is -3.07, while for Chinese investors this coefficient is -1.78. Wage sensitivity for even the most skill-intensive activity (the manufacture of professional, scientific, and controlling equipment) is larger for Foreign investors (-2.10) than for Chinese investors (-1.28), despite the larger Foreign coefficient on the skill-wage interaction.

Greater wage sensitivity by Foreign-funded ventures may reflect a heavier weight placed by these investors on explicit business costs, and less on personal connections, in choosing a location. OECD investors' lack of family and business ties to specific provinces may allow these investors to be more, not less, sensitive to location wage differences. The clustering of overseas-Chinese funded export activities, rather than being evidence of attraction to low-wage havens, as it is often depicted, may instead be explained by an expectation of personal connections to

[21] Grouping of projects into Chinese and Foreign is described by Dean, Lovely, and Wang (2005). The Chinese designation includes those with a partner from Hong Kong, Macao, Taiwan, Malaysia, Indonesia, and the Philippines, with the first three accounting for 87 percent of the total identified with these countries. Projects identified with other sources are denoted Foreign, with the largest shares from the United States and Japan.

protect and promote business interests.[22] Additionally, given previous evidence on differences in skill intensity, it is likely that the activities of Chinese investors in any ISIC category are less skill intensive than that of their Foreign counterparts.[23] This difference may explain why the response of investors using China as an export platform is not as sensitive to wages and skill intensity as Foreign investors investing to serve the local market.

Allowing for Differential Response by Export Market Competition

Race-to-the bottom scenarios typically argue that wage pressures in poor host countries are driven by competition among the low-wage countries to be platforms for exports to richer markets. For example, Ross and Chan (2002, p.10) write, "Wages have fallen as a result of intensified competition to attract factories that sell to the North's markets." We incorporate this hypothesis into our analysis by positing that the price elasticity faced by Chinese exports is a function of the degree of low-wage country competition in the US market. If many rival countries are able to export similar goods to rich markets, subcontractors will have little ability to pass along wage increases through higher product prices. Moreover, if Chinese investors are more intensely engaged in activities that use China as an export platform, we would expect these wage pressures to be more intense for them than they are for Foreign investors. Indeed, Ross and Chan specifically address the wage pressures exerted on Chinese workers by South Korean, Taiwanese, and Hong Kong firms that subcontract with brand-name corporations to do labor-intensive manufacturing.

To explore the hypothesis that competition for Northern markets influences firms' wage

[22] Such expectations are supported by extensive interviews conducted by and summarized in Wang (2001).
[23] Additional support for this possibility comes from Schott (2004), who finds that the unit values of US imports sourced simultaneously from low, medium, and high wage countries are related to country endowments, suggesting that there are important quality differences not captured by standard classifications. Pham-si (2004) provides evidence that the source of much of the growth in these simultaneously sourced exports is China. These studies suggest that even though Chinese EJVs are operating in highly skill-intensive industries, the activities performed in China are not as intensive as the US classification would suggest.

sensitivity, we group industries by the extent to which they compete with other low-wage countries in their exports to the United States. Using US import data, we calculate the shares of imports from low-income countries sourced from China for each 3-digit ISIC industries in 1996.[24] Low-income countries are defined as countries whose GDP per capita (PPP) in 1996 is less than US $5000. The import shares by industries are shown in Table 9. Following the logic of Proposition 1b, we expect the coefficient on an interaction of the wage and this import share measure to be positive – a larger US market share implies less intense competition from other low-wage countries and, hence, lower price elasticity and, thus, less pressure on wage costs. We expect that this effect will be larger for Chinese investors than for Foreign investors. To isolate the effect of market competition from factor intensity, we control for the skill intensity of the industry while exploring this hypothesis.

Table 8 provides the results of the conditional logit analysis, including interactions between the wage and skill intensity and between the wage and US import share of the project's industry. For the control function, we include in the conditional logit the residual of the first-stage IV wage regression and its interaction with skill and share, when these are also interacted with wage. Looking at the first panel of Table 8, we see that for the full sample the wage and its interaction with US import-market share have the predicted signs and are highly significant. The positive sign on the share interaction is evidence that the sensitivity of a firm to local wages is tempered by the degree to which it competes with other exporters.

The second panel indicates that this relationship is not simply a function of skill intensity. While one might wonder if it is the labor-intensive firms that have low market shares and hence greater wage sensitivity, the results in the second panel show that import share has an

[24] We recode the import data from the harmonized system (HS) to ISIC Rev2 using the concordance found at http://www.macalester.edu/research/economics/PAGE/HAVEMAN/Trade.Resources/Concordances/FromHS/hs_isic.txt.

independent effect on wage sensitivity. Controlling for skill intensity, the estimated coefficient for the wage-share interaction is 1.14 and it is highly significant. The wage-skill interaction remains positive and precisely estimated. Although the two estimated interaction coefficients are positive, the estimated wage effect remains negative for every value of industrial import share and factor intensity.

The third panel provides results estimated using only Chinese invested ventures. We see that for this group of projects, all estimated coefficients have the expected signs, but only the wage and its interaction with import share are significant. Comparing these results to those in Table 6, we see that including the wage-share interaction has no effect on the estimated coefficients for the remaining control variables. The estimated coefficient on the wage-share interaction is 1.65, which implies that the wage effect in the industry with the lowest US import share is almost twice as large in absolute value as the effect for firms in an industry with a dominant US market position. This result is consistent with descriptions of Chinese funded ventures as labor-intensive export platforms whose behavior is influenced by their position in the international division of labor. The interaction of wage and skill intensity is not significant for this investor group, although the estimated coefficient remains positive. As before, this evidence is consistent with the view that these ventures do not reflect the factor intensity of the US industry as a whole, but rather subcontract to provide the most labor-intensive activity.

The fourth panel provides similar estimates for the foreign-funded ventures. For this group, the wage-skill interaction is positive and highly significant. Interestingly, however, the estimated coefficient on the wage-share interaction, while of the expected sign, is small in magnitude and not statistically significant. Indeed, adding the wage-share interaction to the equation has very little effect on any other estimated coefficient, as seen by a comparison of the last panels of Table

6 and 8. Controlling for export market share has virtually no effect on the relationship between wage and skill intensity – the estimated coefficient on wage-skill interaction falls in absolute value only by about 3 percent.

We are aware of no other econometric analysis of how export competition influences the wage pressure exerted by foreign investors. Our results support the view that competition for Northern markets influences the weight placed on local wages by investors, but suggests that this link is operative only for those investors who use the host country as an export platform. While we cannot control directly for the share of output that investors expect to export, the identification of source country with export activity allows a window into a previously unexamined link between FDI location choice and competition for rich country markets. We find that Chinese investors are sensitive to their role in the US market when locating their manufacturing facilities on the mainland and that import share is a much more important determinant of wage sensitivity than the skill intensity of the industry.

This contrasts, of course, with the results we estimate for the OECD investors. The weight placed on local wages by these investors appears to be virtually independent of the Chinese export position. Instead, wage pressures for this group are related to skill intensity; more unskilled-labor-intensive industries place a significantly higher weight on wages in choosing a location.

Despite these differences in behavior, which are consistent with earlier characterizations of the two investor groups, OECD investors are more wage sensitive than investors from ethnically Chinese economies. We summarize these differences by calculating the average estimated own-wage elasticities for each industry, for both the Chinese and the Foreign samples, as shown in Table 9. Comparing the last two columns, we see that the Foreign elasticity is larger than the Chinese elasticity for every industry. Some interesting differences emerge when we look across industries. Among Chinese investors, those with joint ventures in petroleum refining and non-ferrous metals have the largest estimated wage elasticities, due to their low US market share, followed closely by food, apparel, and textiles. The smallest elasticities for this group are estimated for miscellaneous petroleum and coal production and for printing. Among Foreign investors, those with ventures in textiles, apparel, wood, footwear, iron and steel, and leather have the largest estimated elasticities. The smallest elasticities for this group are estimated for the manufacture of professional, scientific, and controlling equipment and for printing. These differences across industries suggest that the wage pressure to which an individual worker is subject depends on his or her industry-specific skills.

VI. How Wages Matter and to Whom

Previous microdata studies of firm location choice have found little support for the hypothesis that firms are sensitive to local wages in choosing a local host for their investment. We explore the possibility that omitted variable bias explains the failure to find a deterrent effect for relatively high wages. We introduce to the location-choice context a control-function approach developed by Petrin and Train (2004). Using data from 2884 manufacturing equity-joint-venture projects in China during 1993-1996, we find that standard conditional logit techniques underestimate the sensitivity of investors to local wages and that coefficients estimated using the control function are more than twice as large in absolute value. Although the Chinese circumstances are unique, especially in that they permit the use of the SOE wage as an instrument in our first-stage regression, we see promise in the use of this control-function approach in other contexts.

A second contribution of the paper is new evidence on the nature of firms' attraction to low wages. Using the control-function approach, we find that firms are sensitive to wages in locating their investments. There are significant differences among firms, however, consistent with domestic market orientation. Wage sensitivity of Chinese investors, who are characterized as using China as an export platform, depends on the industry's position in export markets. Firms facing the greatest competition from other low-wage countries are the most sensitive to wages when making location decisions. Wage sensitivity of OECD investors, who are characterized as serving the Chinese domestic market, depends on the skill intensity of the industry. Unskilled-labor-intensive activities are the most sensitive to wages in choosing a host. These results are consistent with the microeconomic theory of the firm and with observations in the literature and the popular press. They direct our attention to the role played in the development process by particular types of activities and they suggest that wage pressures on local hosts change as the development process matures.

Finally, our estimates indicate that OECD investors are more sensitive to wages than investors from ethnically-Chinese economies. For OECD investors in the most labor-intensive industries, the own-wage elasticity of the probability of locating in a province exceeds 2. The clustering of Chinese firms, perhaps due to political, family, or business connections, does not appear to be a response to low wages. Observed outcomes reflect complex calculations by firms faced with local differences on many dimensions that influence business costs.

Appendix: Application of the Control Function Approach

A maintained primitive of the control function approach is that wages are additively separable in the observed (X_j and Z_j) and the unobserved factors (ξ_j), that is, the unobserved factors are mean independent of the observed factors. This assumption implies uncorrelatedness of unobservables and covariates. It enables use of linear regression in the first stage and ensures the consistent estimation of the residual from the first stage. This assumption also suggests a direction for the bias in the wage coefficient estimated without correction. The control function will be positively correlated with EJV location choices and the coefficient on wages will be underestimated in absolute value using standard conditional logit analysis. Thus, elasticities estimated using the control function should be larger than those estimated without it.

When a control function that includes predicted values is added to the estimation, the coefficients are consistent but the standard errors are incorrect. Petrin and Train (2004) use bootstrapping to correct standard errors in their applications. In the first stage, for each bootstrapped sample, we regress the private wage on the exogenous variables and the instrumental variable, SOE wage, for years 1990-1996.[25] The control function in the second stage is a function of the first stage residual (and the interactions of the residual with other covariates when we use interactions of these covariates with wages). We run the conditional logit with this control function and repeat this process 100 times. The variances of these bootstrapped coefficients in the second stage are added to the traditional variance estimates from the conditional logit regression with the control function.[26] We experiment with different orders of

[25] We do not use years after 1996 in the first stage to avoid possible structural changes in wage structure after 1996 due to SOE reforms. We also do not use years before 1990 for similar concerns. Years after 1989 and before 1993 are kept to increase the sample size and the reliability of bootstrapping. However, the direction of bias is consistent when we experiment with different years in the first stage.

[26] Karaca-Mandic and Train (2003) propose alternative standard error correction procedures, but find results very similar to bootstrapping.

the polynomial of the residuals to specify the control function, but only report results with the first-order term because typically higher orders are insignificant and have only a small effect on our bias estimates.

Table A1. First Stage OLS Regression: Dependent Variable is Private Wage

Variables	Coef.	Robust S.E.
Agglomeration †	0.103***	0.01
Local Firms †	-0.018	0.03
Population †	-0.124***	0.04
Skilled Labor Ratio	-0.008***	0.00
Telephone Density †	-0.016	0.03
Road Density †	-0.008	0.02
Private Market Size †	0.108**	0.05
Private Market Size Sqrd †	-0.022***	0.01
Change in State Ownership	-0.284	0.28
SEZ or OCC	-0.077*	0.04
Regional Fixed Effects		
Central	0.061	0.04
Coastal	0.042	0.05
Northeast	-0.005	0.04
Northwest	-0.043	0.04
SOE Wage †	0.840***	0.09
# of Obs	196	
R2	0.86	

Notes: "***", "**" and "*" denote significance levels at 1 percent, 5 percent and 10 percent respectively; variables with "†" are in logarithms; variables are lagged by one year; Gansu and Tibet excluded.

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Table 1: Empirical Studies of Aggregate FDI Flows

Authors	Sample Years	Sample Country	FDI Measure	Wage Measure	Skill Measures	Method	Results
Coughlin and Segev (1999)	1990-1997	China, by province	FDI inflow, in constant US $	Average Provincial Wage	Average Labor Productivity; Share of Population Illiterate	OLS; Spatial Dependence Error Correction	Wage and illiteracy significantly negatively, productivity significantly positively related
Wei, Liu, Parker, Vaidya (1999)	1985-1995	China, by province	Pledged and realized FDI inflow, deflated	Real Provincial Wage, adjusted for productivity (industrial output/emp)	Share of scientists and researchers in total employment	OLS; Error Components Model; First-order autocorrected	Wage and scientist share significantly negatively related in all models
Cheng and Kwan (2000)	1985-1995	China, by province	Stock of FDI	Real Wage	Share of Population by Education Level	GMM; First-differenced GMM; wage treated as endogenous	Real wage significantly negatively related to lagged stock
Fung and Iizaka (2002)	1991-1997	China, by region	FDI inflow from Japan, U.S., Taiwan, and HK	Average lagged nominal wage	Number of students enrolled in secondary and higher education	GLS	Wage significantly negatively related to FDI inflow
Gao (2002)	1996-1999	China, by province	FDI provincial shares from 14 source countries	Average real wage	Shares of employed by education level	OLS, random effects	Wage not significant determinant; sometimes positive coeff.
Fung, Iizaka, and Siu (2003)	1990-2000	China, by region	FDI inflow from Japan and HK	Average lagged nominal wage	Share of population enrolled in higher education	GLS	Wage significantly negatively related to FDI inflow

Table 2: Empirical Studies of Foreign Affiliate Location Choice

Authors	Sample Years	Sample Country	FDI Measure	Wage Measure	Skill Measures	Method	Results
Ondrich and Wasylenko (1993)	1978-1987	United States, by state	New foreign-owned plants	Real avg. hourly production earnings	None	Conditional logit; nested logit	Wage negative but not significant determinant
Head and Ries (1996)	1984-1991	China, 54 cities	New equity joint ventures	Industrial wage bill/ industrial workers	Value of industrial output/ industrial workers	Conditional logit	Wage not significant, sometimes positive coeff.
Devereux and Griffith (1998)	1980-1994	European counries, by US companies	US foreign affiliates	Industry unit labor costs	None	Nested logit	Wage not significant, sometimes positive coeff.
Head, Ries, and Swenson (1999)		United States, by state	New Japanese-owned plants	Average manufacturing wage	None	Conditional logit	Wage not significant, sometimes positive coeff.
List and Co (2000)	1986-1993	United States, by state	New foreign-owned plants	Total payroll divided by employees	None	Conditional logit	Wage not significant, sometimes positive coeff.
Keller and Levinson (2002)	1977-1994	United States, by state	New foreign-owned plants	Average hourly production earnings	None	Count data models (negative binomial and alternatives)	Wage not significant, sometimes positive coeff.
Head and Mayer (2004)	1884-1995	European Union, by region	Japanese owned affiliates	Average wage in industry, by region	None	Conditional logit; nested logit	Wage not significant, sometimes positive coeff.
Amiti and Javorcik (2005)	1998-2001	China, by province	Change in number of foreign firms	Change in average provincial wage	None	OLS; NLS;	Wage negative significant determinant

Table 3: Data Definitions and Sources

Variable	Definition	Source	Mean
EJV project: Location Source Industry	 Province Chinese=Macao, Taiwan, Hong Kong, other South Asian countries Foreign=all other countries 3-digit ISIC Rev.2 classification	 *Almanac of China's Foreign Economic Relations and Trade*, various years, Dean, Lovely and Wang (2005)	
SOE Wage	Average wage for industrial workers in state-owned enterprises, in 1990 RMB Yuan/year	Branstetter and Feenstra (2002)	7.93 †
Private Wage	Average wage for industrial workers in other enterprises (private, foreign and etc), in 1990 RMB Yuan/year	Branstetter and Feenstra (2002)	8.05 †
Agglomeration	Cumulative value of real contracted FDI, from 1983 until t-1, in millions of 1980 US dollars	Coughlin, et al. (2000)	12.95 †
Local Firms	Number of SOE and collective industrial enterprises at the township level and above, in thousands	*China Statistical Yearbook*, various years	9.36 †
Population	Province population, in millions	*China Statistical Yearbook*, various years	3.44 †
Skilled Labor Ratio	Percent of population who have a senior secondary school education level or above	*China Statistical Yearbook*, various years and calculations by authors	12.08
Telephone Density	Number of Urban telephone subscribers per 1000 persons	*China Statistical Yearbook*, various years	9.97 †
Road Density	Road (km)/land area (km^2)	*China Statistical Yearbook*, various years	5.27 †
Private Market Size	Real Provincial GDP*(1–SOE share), where SOE share is the production share of SOEs; GDP is value in billions of 1990Yuan	*China Statistical Yearbook*, various years	3.47 †
Change in State Ownership	Difference between shares of industrial output from SOEs in year t and t-1	*China Statistical Yearbook*, various years	-0.04
SEZ or OCC	Dummy variable for a province with SEZ or Open Coastal City	Constructed by authors	0.43
Skill Intensity	Number of non-production workers divided by number of production workers, by ISIC 3-digit classification	NBER-CES Manufacturing Industry Database; concorded by authors	0.45
Import Share	Chinese share of U.S. imports from low income countries in 1996, by ISIC 3-digit classification	U.S. Import and Export Data, www.internationaldata.org; concorded by authors	0.62

Note: Variables with "†" are in logarithms.

Table 4: Provincial Characteristics, Period Averages (1993-1996)

	Annual Wage (1990 yuan)	Cum FDI (million 1980 USD)	# of local firms (000s)	POP (mils)	Share of Skilled Worker	Tele-phone per 1000	Road (km per km2)	Output Share of SOEs	Private Market Size (billion yuan)
Anhui	3083	353	23	59	7	13	0.23	0.41	53
Beijing	4695	1981	7	11	32	119	0.71	0.51	34
Fujian	3561	4499	12	32	8	28	0.36	0.23	73
Guangdong	4970	13876	25	67	11	47	0.38	0.25	192
Guangxi	3045	933	11	45	8	12	0.17	0.50	37
Guizhou	2810	87	6	34	6	6	0.18	0.71	10
Hainan	4476	1336	1	7	12	26	0.39	0.53	9
Hebei	2701	566	21	64	8	18	0.26	0.38	82
Heilongjiang	2819	434	17	37	15	30	0.10	0.72	28
Henan	2426	450	20	90	8	10	0.29	0.40	83
Hubei	2574	704	23	57	10	17	0.26	0.49	58
Hunan	3346	475	23	63	9	15	0.27	0.48	54
In. Mongolia	2122	78	9	22	13	22	0.04	0.68	13
Jiangsu	3489	4273	39	70	12	27	0.25	0.23	184
Jiangxi	2565	293	16	40	8	12	0.20	0.50	29
Jilin	2553	333	13	26	17	33	0.15	0.65	20
Liaoning	3390	2064	26	41	14	37	0.29	0.49	75
Ningxia	2462	11	2	5	11	22	0.16	0.74	2
Qinghai	2850	5	1	5	11	17	0.02	0.83	1
Shaanxi	3042	483	13	35	12	14	0.18	0.61	20
Shandong	2691	2929	25	87	9	16	0.31	0.30	160
Shanghai	5654	3514	9	14	29	124	0.56	0.46	65
Shanxi	2876	107	11	30	12	17	0.20	0.49	27
Sichuan	2960	759	37	112	7	10	0.18	0.44	94
Tianjin	4213	1039	8	9	22	65	0.34	0.41	26
Xinjiang	3027	64	6	16	14	20	0.02	0.71	12
Yunrjan	3021	113	7	39	5	11	0.17	0.73	16
Zhejiang	3684	1251	36	43	9	32	0.33	0.19	127

Table 5: Conditional Logit Analysis of EJV Provincial Location Choice

Variables	Full Sample No Regional Fixed Effects		Full Sample Regional Fixed Effects		Full Sample First-order Control Function		Full Sample Second-order Control Function	
	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.
Private Wage †	-1.70***	0.17	-0.96***	0.20	-2.12***	0.47	-2.19***	0.49
Agglomeration †	0.18***	0.05	0.32***	0.05	0.46***	0.08	0.49***	0.08
Local Firms †	0.65***	0.10	1.12***	0.12	1.02***	0.13	1.03***	0.14
Population †	1.00***	0.15	1.75***	0.16	1.66***	0.19	1.58***	0.19
Skilled Labor Ratio	0.08***	0.01	0.12***	0.01	0.10***	0.01	0.10***	0.01
Telephone Density †	0.57***	0.10	0.32***	0.11	0.55***	0.15	0.49***	0.15
Road Density †	0.39***	0.08	0.12	0.10	0.14	0.11	0.14	0.11
Private Market Size †	-1.72***	0.23	-3.04***	0.27	-3.01***	0.29	-2.94***	0.30
Private Market Size Sqrd †	0.20***	0.02	0.21***	0.02	0.20***	0.03	0.20***	0.03
Change in State Ownership	-2.77***	0.78	-5.51***	0.83	-6.07***	1.06	-6.22***	1.04
SEZ or OCC	0.81***	0.10	1.28***	0.15	1.11***	0.18	1.14***	0.18
Regional Fixed Effects								
Central			1.55***	0.17	1.52***	0.18	1.58***	0.18
Coastal			1.74***	0.17	1.71***	0.20	1.63***	0.20
Northeast			1.01***	0.16	0.70***	0.21	0.71***	0.22
Northwest			0.38*	0.22	0.21	0.23	0.24	0.23
Residual					1.77***	0.61	1.70***	0.65
Residual^2							-6.94*	3.58
# of Obs	80752		80752		80752		80752	
LR test	3335		3497		3513		3526	
Likelihood	-7943		-7862		-7854		-7847	
Pseudo R2	0.17		0.18		0.18		0.18	

Notes:

1. All covariates are lagged by one year;
2. Variables with "†" are in logarithms;
3. "***", "**" and "*" denote significance levels at 1 percent, 5 percent and 10 percent respectively;
4. Gansu and Tibet are excluded as no EJVs located there during 1993-1996.

Table 6: Estimated Own and Cross Wage Elasticities, by Province

Provinces	Without Control Function			With Control Function		
	Predicted Probability	Own Elasticity	Cross Elasticity	Predicted Probability	Own Elasticity	Cross Elasticity
Anhui	0.018	-0.940	0.017	0.017	-2.087	0.036
Beijing	0.076	-0.884	0.073	0.081	-1.950	0.173
Fujian	0.027	-0.932	0.025	0.028	-2.064	0.059
Guangdong	0.134	-0.829	0.128	0.128	-1.851	0.272
Guangxi	0.011	-0.946	0.011	0.010	-2.102	0.021
Guizhou	0.002	-0.955	0.002	0.002	-2.118	0.005
Hainan	0.005	-0.952	0.005	0.007	-2.109	0.014
Hebei	0.062	-0.898	0.059	0.062	-1.991	0.132
Heilongjiang	0.026	-0.932	0.025	0.026	-2.068	0.055
Henan	0.023	-0.935	0.022	0.024	-2.072	0.051
Hubei	0.035	-0.923	0.033	0.034	-2.051	0.072
Hunan	0.025	-0.933	0.024	0.026	-2.068	0.055
Inner Mongolia	0.005	-0.953	0.004	0.004	-2.114	0.009
Jiangsu	0.181	-0.784	0.173	0.190	-1.720	0.403
Jiangxi	0.015	-0.943	0.014	0.015	-2.091	0.032
Jilin	0.021	-0.937	0.020	0.019	-2.082	0.041
Liaoning	0.059	-0.901	0.056	0.060	-1.995	0.128
Ningxia	0.002	-0.955	0.002	0.002	-2.119	0.004
Qinghai	0.001	-0.956	0.001	0.001	-2.120	0.003
Shaanxi	0.008	-0.949	0.008	0.008	-2.106	0.017
Shandong	0.118	-0.844	0.113	0.114	-1.881	0.242
Shanghai	0.049	-0.910	0.047	0.044	-2.029	0.094
Shanxi	0.008	-0.950	0.007	0.007	-2.107	0.016
Sichuan	0.010	-0.947	0.010	0.011	-2.099	0.024
Tianjin	0.026	-0.932	0.025	0.025	-2.070	0.053
Xinjiang	0.001	-0.956	0.001	0.001	-2.121	0.002
Yunnan	0.001	-0.956	0.001	0.001	-2.120	0.003
Zhejiang	0.053	-0.906	0.050	0.052	-2.014	0.109

Notes:

1. The left panel is associated with the second column of Table 5;
2. The right panel is associated with the third column of Table 5.

Table 7: Conditional Logit Analysis of EJV Provincial Location Choice, Skill-Intensity Interactions, Various Samples

Variables	Full Sample		Full Sample Control Function		Chinese Sample Control Function		Foreign Sample Control Function	
	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.
Private Wage †	-1.43***	0.24	-2.66***	0.55	-2.15***	0.61	-3.81***	0.72
Wage†*Skill Intensity	1.07***	0.33	1.13***	0.34	0.84**	0.42	1.67***	0.55
Agglomeration †	0.32***	0.05	0.47***	0.08	0.53***	0.09	0.31***	0.12
Local Firms †	1.12***	0.12	1.02***	0.13	0.88***	0.16	1.29***	0.21
Population †	1.75***	0.16	1.65***	0.20	1.79***	0.23	1.39***	0.28
Skilled Labor Ratio	0.12***	0.01	0.10***	0.01	0.07***	0.01	0.15***	0.02
Telephone Density †	0.31***	0.11	0.55***	0.16	0.61***	0.20	0.52**	0.20
Road Density †	0.12	0.10	0.14	0.11	0.13	0.14	0.20	0.16
Private Market Size †	-3.04***	0.27	-3.01***	0.29	-2.76***	0.36	-3.33***	0.43
Private Market Size Sqrd †	0.21***	0.02	0.20***	0.03	0.16***	0.04	0.28***	0.04
Change in State Ownership	-5.50***	0.83	-6.08***	1.09	-6.52***	1.29	-6.20***	1.49
SEZ or OCC	1.28***	0.15	1.11***	0.17	0.75***	0.22	1.51***	0.27
Regional Fixed Effects								
Central	1.55***	0.17	1.52***	0.17	1.57***	0.21	1.23***	0.28
Coastal	1.74***	0.17	1.71***	0.19	1.94***	0.23	1.30***	0.30
Northeast	1.01***	0.16	0.69***	0.22	0.63**	0.26	0.64**	0.31
Northwest	0.38*	0.22	0.20	0.25	0.00	0.30	0.35	0.36
Residual			1.81***	0.70	1.60***	0.72	2.03**	0.93
# of Obs	80752		80752		47908		32844	
LR test	3507		3524		2081		1598	
Likelihood	-7856		-7848		-4661		-3110	
Pseudo R2	0.18		0.18		0.18		0.20	

Notes:

1. All covariates are lagged by one year;
2. Variables with "†" are in logarithms;
3. "***", "**" and "*" denote significance levels at 1 percent, 5 percent and 10 percent respectively;
4. The interactions of the predicted residual and skill intensity are insignificant, and hence are not included.
5. Gansu and Tibet are excluded as no EJVs located there during 1993-1996.

Table 8: Conditional Logit Analysis of EJV Provincial Location Choice, Using Skill-Intensity and Import-Share Interactions, Various Samples

Variables	Full Sample Control Function		Full Sample Control Function		Chinese Sample Control Function		Foreign Sample Control Function	
	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.
Private Wage †	-2.89***	0.54	-3.22***	0.55	-3.07***	0.67	-3.90***	0.77
Wage†*Share	1.24***	0.33	1.14***	0.34	1.65***	0.43	0.18	0.47
Wage†*Skill Intensity			0.83**	0.38	0.63	0.49	1.63***	0.56
Agglomeration †	0.46***	0.08	0.47***	0.08	0.53***	0.09	0.31***	0.12
Local Firms †	1.03***	0.13	1.03***	0.13	0.89***	0.16	1.29***	0.21
Population †	1.66***	0.18	1.65***	0.19	1.78***	0.23	1.39***	0.28
Skilled Labor Ratio	0.10***	0.01	0.10***	0.01	0.07***	0.01	0.15***	0.02
Telephone Density †	0.54***	0.15	0.54***	0.16	0.60***	0.19	0.52**	0.20
Road Density †	0.14	0.11	0.14	0.11	0.13	0.14	0.20	0.16
Private Market Size †	-3.03***	0.29	-3.02***	0.29	-2.78***	0.37	-3.33***	0.44
Private Market Size Sqrd †	0.21***	0.03	0.21***	0.03	0.16***	0.04	0.28***	0.04
Change in State Ownership	-6.11***	0.99	-6.11***	1.08	-6.59***	1.27	-6.20***	1.45
SEZ or OCC	1.11***	0.18	1.11***	0.19	0.75***	0.22	1.51***	0.26
Regional Fixed Effects								
Central	1.52***	0.18	1.52***	0.18	1.57***	0.21	1.23***	0.28
Coastal	1.72***	0.19	1.72***	0.20	1.95***	0.24	1.30***	0.30
Northeast	0.70***	0.21	0.69***	0.22	0.64**	0.26	0.64**	0.31
Northwest	0.21	0.23	0.21	0.23	0.01	0.30	0.35	0.37
Residual	2.92***	0.86	2.48***	0.93	2.95**	1.15	2.03**	0.97
Residual*Share	-1.87**	0.92	-2.02**	0.98	-2.54**	1.21		
Residual*Skill Intensity			1.29	1.23	0.49	1.53		
# of Obs	80752		80752		47908		32844	
LR test	3528		3539		2097		1598	
Likelihood	-7846		-7841		-4653		-3110	
Pseudo R2	0.18		0.18		0.18		0.20	

Notes:

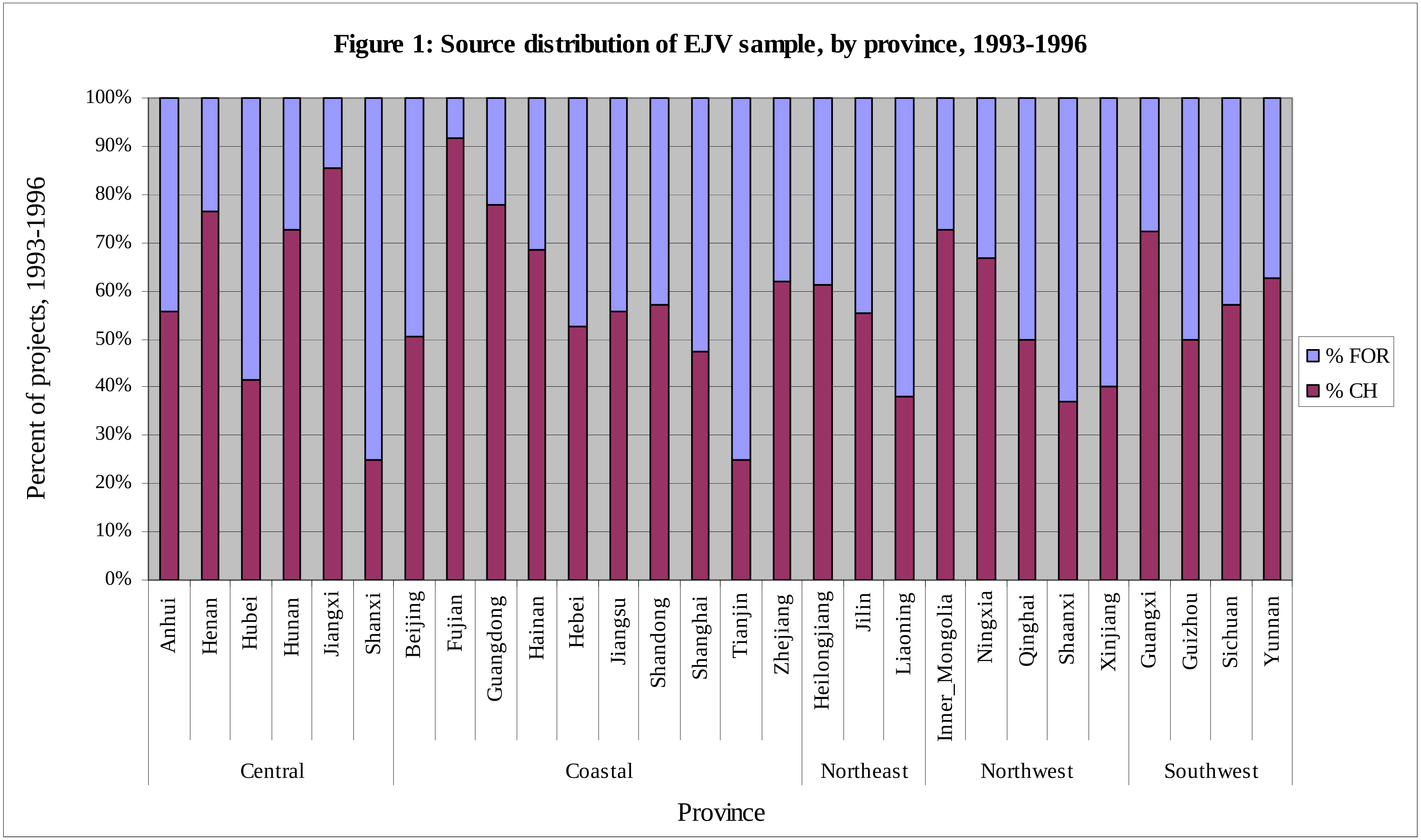
1. All covariates are lagged by one year;
2. Variables with “†” are in logarithms;
3. “***”, “**” and “*” denote significance levels at 1 percent, 5 percent and 10 percent respectively;
4. Gansu and Tibet are excluded as no EJVs located there during 1993-1996.

Table 9: Average Estimated Own Wage Elasticity, by Industry

ISIC	Industry Name	Skill Ratio	Import Share	Full Sample	Chinese Sample	Foreign Sample
353	Petroleum Refineries	0.56	0.01	-2.65	-2.61	-2.88
372	Non-ferrous Metals	0.33	0.18	-2.64	-2.47	-3.21
322	Apparel	0.19	0.32	-2.60	-2.33	-3.40
311	Food	0.33	0.24	-2.57	-2.37	-3.20
321	Textiles	0.17	0.39	-2.54	-2.23	-3.43
371	Iron and Steel	0.28	0.40	-2.44	-2.16	-3.24
331	Wood	0.21	0.49	-2.40	-2.05	-3.34
355	Rubber	0.30	0.52	-2.29	-1.95	-3.19
332	Furniture	0.26	0.69	-2.14	-1.71	-3.22
313	Beverages	0.97	0.20	-2.11	-2.05	-2.20
323	Leather	0.25	0.72	-2.11	-1.66	-3.23
324	Footwear	0.18	0.79	-2.10	-1.60	-3.34
362	Glass	0.23	0.79	-2.05	-1.56	-3.26
341	Paper	0.31	0.74	-2.04	-1.59	-3.15
369	Mineral	0.36	0.71	-2.04	-1.61	-3.07
351	Industrial Chemicals	0.67	0.50	-2.03	-1.76	-2.62
361	Pottery	0.27	0.81	-2.00	-1.50	-3.20
390	Other	0.43	0.74	-1.95	-1.52	-2.95
381	Fabricated Metal	0.37	0.85	-1.88	-1.39	-3.03
356	Plastic	0.30	0.91	-1.86	-1.33	-3.13
383	Electric Machinery	0.58	0.75	-1.82	-1.42	-2.72
384	Transport	0.64	0.81	-1.71	-1.28	-2.62
382	Non-electric Machinery	0.55	0.89	-1.69	-1.21	-2.74
385	Professional	0.98	0.67	-1.59	-1.30	-2.11
354	Misc. Petroleum and Coal	0.54	1.00	-1.58	-1.04	-2.74
352	Other Chemicals	0.86	0.80	-1.54	-1.17	-2.27
342	Printing	0.90	0.85	-1.46	-1.07	-2.20

Data Sources:
NBER-CES Manufacturing Industry Database: http://www.nber.org/nberces/nbprod96.htm
NBER U.S. Import Data (1972-2001): http://cid.econ.ucdavis.edu/data/sasstata/usiss.html

Note: The last three columns are associated with the last three columns of Table 3 respectively.



Source: Dean, Lovely, and Wang (2005)

Figure 2: Skill intensity of EJV sample, by province, 1993-1996