Several sorting scenarios could also affect the estimate. For example, risk-averse households may move to judicial foreclosure states because they value the implicit insurance; these households may also be *savers* who put down large down payments. Alternatively, borrowers who plan to default on their loans and exploit the judicial protections may move to these states. If lenders cannot identify these households, they may tighten loan terms to compensate for the higher expected losses. These scenarios may not be important empirically, however, if the sorting variables are correlated with control variables in the analysis, or if factors such as proximity to a job, local income tax rates, or school spending matter more than foreclosure laws in the household location decision.

Although it is not possible to give the loan size coefficient a clean interpretation, it seems clear that the mortgage market reaches a different equilibrium in judicial foreclosure states. In these states, borrowers may pay more for their mortgages, purchase smaller houses, or have difficulty becoming homeowners. But borrowers are not necessarily worse off: they may value the insurance provided by the laws. Homeownership might even increase if the judicial protections help borrowers remain in their homes. Although judicial requirements seem to impose costs on borrowers, a full welfare assessment will also require estimates of the law's benefits.

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THE INFLUENCE OF CORRUPTION AND LANGUAGE ON THE PROTRADE EFFECT OF IMMIGRANTS: EVIDENCE FROM THE AMERICAN STATES

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Abstract—The protrade effect of immigrants on the bilateral export performance of the 50 American states and the District of Columbia with respect to 87 foreign countries is studied. This effect, which posits that a greater number of immigrants in a host location leads to increased trade between the host and the immigrants' origin country, has been supported in a number of studies. Here, we extend this approach and find that the immigrant effect is greater when the origin country's political system is more corrupt and less important when Spanish or English is the language of the origin country. State-level export data averaged over the 1990–1992 period are used.

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I. Introduction

THE importance of channels through which individuals are made aware of opportunities for advantageous exchange and the evolution of institutions that provide assurance that agreements will be honored has

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been the focus of papers by Avner Greif, James Rauch, and others.[1] Presumably, the needed information and trust are more difficult to obtain when exchange occurs at a distance or crosses linguistic or cultural boundaries, as is the case with much of international trade. The buyer needs assurance that the merchandise is as claimed, and the supplier must be confident in receiving the agreed-upon payment.[2] Finally, for trade to be profitable, merchandise must move through the distribution channels in a timely fashion, and the costs, including bribes, needed to move the merchandise must be anticipated with some certainty.

Migrant networks are capable of providing this information on opportunities and on the reputations of potential trading partners, and of providing sanctions to reduce opportunistic cheating on agreements. Since Gould (1994), who found an important protrade role for the foreign-born in both the import and export trade of the United States with their countries of origin, the protrade effect of immigrants has been confirmed for Canada by Head and Ries (1998), for the United Kingdom by Girma and Yu (2002), and for the United States in the 1870–1910 period by Dunlevy and Hutchinson (1999). Wagner, Head, and Ries (2002) extended the study of the immigrant-trade link using the Canadian province as the unit of domestic observation.

In this paper we use data at the level of the American state to test the influence of the foreign-born on the bilateral exports of their states of residence to their countries of origin.[3] Our focus is not just on the overall protrade effect, but also on how the strength of that effect depends on the level of corruption in the export destination (immigrants' origin) country and on the similarity of language and institutions between the United States and the export destination country.

Data from the World Institute for Strategic Economic Research (WISER) on average annual exports of manufactured goods over the 1990–1992 period, by state and by country of destination, are combined with data on the number of foreign-born, by state of residence in 1990 and by country of birth, in a gravity model to test the protrade immigrant hypothesis. The model is then augmented to test several conjectures. Notably, we find that the protrade role of immigrants is enhanced if their origin country's political system is more corrupt and weakened if their native language is English or Spanish.

II. The Importance of Trust and Information

Our starting point is the proposition that immigrants in the host country share strong ethnic ties with persons in their origin communities; these networks are a form of social capital that promotes economic contracting and trade.[4] Exports will be greater, ceteris paribus, between U.S. states and destination countries when those countries are sources of larger numbers of immigrants in the particular states. This is what has been found in the studies cited earlier and will be expanded upon here.

These networks are more important when it is more costly to obtain information about the potential trading partner and where it is more difficult to navigate the bureaucratic and commercial environment of the potential partner. This gives rise to several corollaries:

First, immigrant networks are more valuable when destination markets are less transparent or more subject to corruption.[5] The role of immigrants, therefore, will be greater in promoting bilateral trade when the political climate in their origin country is more corrupt. Overall trade with such countries, however, will otherwise be below normal.

Second, immigrant networks are more valuable when the native population in the host country is less able to master the language of the potential trading partner. Here, we assume that members of the immigrant community are more likely than natives in the host country to be competent in both their host-country and origin-country languages.[6]

Third, Girma and Yu have proposed that information possessed by immigrants is less useful when the immigrants are from a country whose institutions are similar to those of the host country. Therefore, the more similar the institutions of the immigrant source (export destination) country are to those of the host country, the weaker will be the protrade effect of the immigrants.[7]

III. Model and Data

The basic gravity equation relates exports between i and j as an increasing function of the incomes and populations of the two trading entities and as a decreasing function of the distance between them; factors that alter the costs and benefits of trade are then added to the list of regressors on the basis of any of a variety of possible underlying theoretical derivations.[8] We enter migrant stock (MS) into the model to reflect the sought-after information and cultural network effects. The model to be estimated is

$$\begin{aligned}\ln EXPORTS_{ij} = f(&\ln MS_{ij}, \ln PCGSP_i, \ln GSP_i,\\ &\ln PCGDP_j, \ln GDP_j, \ln DISTANCE_{ij}, Z_{ij}),\end{aligned} \tag{1}$$

where

- $\ln EXPORTS_{ij}$ denotes (the logarithm of) the dollar value of exports of manufactures from state i to country j averaged over 1990 through 1992,
- MS_{ij} denotes the number of persons born in country j residing in state i as enumerated in the 1990 Census,
- GSP_i and GDP_j denote, respectively, the gross state product of state i and the gross domestic product of country j in U.S. dollars in 1990,
- $PCGSP_i$ and $PCGDP_j$ denote, respectively, the per capita GSP of state i and the per capita GDP of country j in 1990,
- $DISTANCE_{ij}$ denotes the distance from the principal or most central city of state i to the capital or major city of country j,
- Z_{ij} denotes other variables used to augment this standard form that allow us to test the corollaries advanced in the previous section.

[1] Rauch (1999), Rauch and Trindade (2002). See Greif (2000) for a general reference.

[2] See Marcouiller (2001) and Anderson and Marcouiller (2002) for further discussion.

[3] Coughlin and Wall (2003) provide a survey of the literature on state-level exports.

[4] Granovetter's (1973) discussion or *strong* and *weak* ties underlies our network theory. A fuller discussion of the role of strong and weak ties is in an extended version of this paper available from the author or online at http://www.sba.muohio.edu/dunlevja/Research.htm.

[5] Anderson and Marcouiller (2002, p. 342) cite a World Bank survey that lists corruption as a significant obstacle to business worldwide.

[6] Hutchinson (2002) applies the *language distance* measure of Chiswick and Miller (1998) to international trade but does not incorporate the role of immigrants. Language distance does not play a role in the present paper.

[7] A fourth corollary, that quality information and trust are more important for trade in differentiated goods than for homogeneous products and that the immigrant effect should therefore be more pronounced in the case of more highly differentiated goods, has been confirmed by Gould (1994), Rauch (1999), Rauch and Trindade (2002), and Dunlevy and Hutchinson (1999), among others. A fifth corollary allows that overseas stays by the native-born from the host country may also be trade-creating. These issues are addressed in the extended version of the paper, with inconclusive results.

[8] Basic references include Bergstrand (1985).

TABLE 1.—TOBIT ESTIMATION WITH CORRUPTION, LANGUAGE, AND INSTITUTIONAL SIMILARITY

	Without Fixed Effects		With Fixed Effects		
	(1)	(2)	(3)	(4)	(5)
ln MS_{ij}	0.29 (9.54)	0.37 (11.03)	0.47 (14.54)	0.24 (7.22)	0.39 (10.96)
$CORRUPTION_j \times \ln MS_{ij}$		0.17 (14.33)	0.15 (13.13)		
$CORRUPTION_j$		−0.93 (13.31)	−0.87 (12.90)		
$ENGLISH_j \times \ln MS_{ij}$		−0.17 (3.41)	−0.17 (3.90)		
$ENGLISH_j$		2.12 (7.64)	2.06 (8.24)		
$SPANISH_j \times \ln MS_{ij}$		−0.44 (9.80)	−0.42 (10.37)		
$SPANISH_j$		3.57 (13.64)	3.70 (15.72)		
$INSTITUTIONAL\ SIMILARITY_j \times \ln MS_{ij}$		−0.05 (0.50)	0.09 (1.05)		
$INSTITUTIONAL\ SIMILARITY_j$		0.13 (0.20)	−0.71 (1.21)		
ln GSP_i	1.82 (30.39)	1.82 (36.86)		1.83 (30.44)	
ln per capita GSP_i	−1.34 (7.77)	−1.26 (7.97)		−1.26 (8.59)	
ln GDP_j	3.09 (45.80)	3.19 (34.92)	3.01 (35.69)		
ln per capita GDP_j	−0.94 (28.47)	−0.95 (25.29)	−0.84 (23.89)		
$DISTANCE_j$	−1.40 (15.34)	−1.02 (9.52)	−0.52 (5.11)	−3.21 (15.86)	−1.90 (8.75)
Pseudo R^2	0.17	0.19	0.23	0.23	0.26

Note: Absolute *t*-values are reported in parentheses; unconditional marginal effects are reported.

Given the geographically disaggregated nature of our data, the exports from a number of states to given foreign destinations are zero in value. This requires two adjustments: first, the logarithm of $EXPORTS_{ij}$ is redefined as $\ln(EXPORTS_{ij} + 1)$; second, because the resulting value of $\ln(EXPORTS_{ij} + 1)$ is 0 in those cases where $EXPORTS_{ij}$ is 0, the model is estimated using the tobit procedure.

The data sources are relatively standard, so we will discuss further only the export data.[9] These are from the origin of movement series collected by the U.S. Census Bureau and released by WISER. These data, available at the two-digit SIC level, report the point of origin from which exports begin their journey. *Point of origin* generally refers to the state in which the factory that produced the item is located, or the location of a distributor, warehouse, or cargo-processing facility. Although the state of origin of movement and the state of production are not always identical, for manufactured goods they usually are.[10] We limit ourselves to exports of manufactures.

Our data comprise exports of the 50 U.S. states and the District of Columbia, referred to as *states,* to 87 foreign countries.[11,12] The destination countries span all continents and all ranges of economic development; some are major contributors of immigrants to the United States, others contribute few immigrants; some are major markets for U.S. exports, others are minor trading partners.

IV. Empirical Findings

Estimation of the basic gravity model with *MS* is reported in column 1 of table 1. The estimated marginal effects on observed exports, unconditional on whether the observation on exports is censored, are reported.[13] The estimated elasticity of *MS* on *EXPORTS* is a statistically and economically significant 0.29; this is some two times as large as the effect reported by Head and Ries (1998), by Girma and Yu (2002), and by Wagner, Head, and Ries (2002). The proposition offered above is affirmed.

All coefficients on the standard gravity variables are statistically different from 0. The elasticities on GSP (1.82) and on foreign GDP (3.09) are in the upper range of what is usually reported. Though the negative coefficients obtained on the per capita income coefficients are counterintuitive they are consistent with what Gould found for exports from the United States as a whole.[14] The estimated elasticity of distance, −1.40, is standard.

The model is modified to address the influence of corruption in the export destination country, the roles of both English and Spanish languages, and the effect of institutional similarity. These results are reported in column 2.

Here, *MS* obtains a statistically significant coefficient of 0.37, somewhat stronger than what was obtained in the baseline specification. The estimated coefficients on the standard gravity variables are little affected by the inclusion of corruption, language, and institutional similarity.

The variable *CORRUPTION* is expressed in terms of deviations from its own mean, and it is configured so that a higher value is associated with greater corruption.[15] It appears in the regression in two ways. First, it is interacted with the logarithm of *MS* in order to determine if the protrade effect of immigrants is greater when the export destination country has a higher level of corruption. Second, it is added as a regressor to capture any intercept-shift effects of corruption on trade flows.

The estimated coefficient on the interaction of *CORRUPTION* and *MS* is a statistically significant 0.17, implying, as hypothesized, that the protrade role for immigrants increases if their home country's political environment is more corrupt. For instance, when combined with the estimated *MS* elasticity of 0.37, the implicit protrade elasticity of immigrants from Guatemala, whose modified corruption index

[9] Data sources are given in the appendix.

[10] See the U.S. Department of Commerce, International Trade Administration, "State Exports."

[11] Of the 4,437 observations, 239 have an export value of 0. Summary statistics on export volume and migrant stock are available online and from the author.

[12] The cross-sectional nature of our data prevents a direct test of the hypothesis that the direction of causality runs from migrants to trade. Gould (1994, p. 310, footnote 17) gives reasons for believing that cross-section estimation will, in fact, capture the assumed effects. Dunlevy and Hutchinson (1999) provide evidence supporting the asserted direction of causation; they also present results consistent only with causation running from migration to trade.

[13] The dtobit routine of the Stata statistical package was used to obtain the estimates.

[14] Gould uses the logarithms of population and GDP as regressors and obtains a positive coefficient of 4.10 on U.S. population. Because of the linear relationship among the logarithms of population, GDP, and GDP per capita, this implies an elasticity of −4.10 on American per capita GDP.

[15] Observations on *CORRUPTION* do not exist for Cape Verde Islands, Grenada, Laos, or Western Samoa. As a result, 204 observations are lost in those estimations that have *CORRUPTION* among their explanatory variables. For this subset, 174 of the 4,233 observations have a value of 0 for the dependent variable.

is 1.60, is 0.64, whereas the elasticity for immigrants from Hong Kong, whose modified corruption index is -1.40, is 0.13. The estimated coefficient on the *CORRUPTION* intercept shift variable is a statistically significant -0.93, which supports the second part of the first corollary that, ceteris paribus, trade is deterred by corruption. Whereas this latter result is well grounded in the literature, we believe the finding regarding the interaction of the role of immigrant networks and corruption on trade is novel.

Similarity of language between two countries has been found to be trade-creating.[16] The second corollary, however, asserts that similarity of language should reduce the contribution of immigrants to trade promotion. The third corollary similarly argues that immigrants from countries with legal and commercial institutions similar to those of the United States will contribute less to trade creation than immigrants from other countries. We consider the effect of language and institutions by entering variables that are the interaction of *ENGLISH, SPANISH,* and *INSTITUTIONAL SIMILARITY* with the logarithm of *MS*. We also expect that trade with these countries will involve lower transactions costs, and that exports to them, ceteris paribus, will be greater. This effect is tested using intercept-shift dummy variables.[17]

Both language–migrant-stock variables are statistically significant: -0.17 for the *ENGLISH* $\times$ *MS* interaction and -0.44 for the *SPANISH* $\times$ *MS* interaction. This suggests that the protrade effect of migrants from English-speaking countries is almost 50% weaker than that for the reference group; migrants from Spanish-speaking countries are estimated to have no protrade effect whatsoever. The coefficients of the intercept-shift variables are both statistically significant—2.12 for English-speaking countries and 3.57 for Spanish-speaking countries—indicating that exports to English-speaking and to Spanish-speaking countries, ceteris paribus, are well above the norm. The second corollary receives strong support. Neither institutional similarity coefficient is meaningfully different from 0; the third corollary is not supported.

One might argue that the modified gravity model fails to control for heterogeneity across the American states and across the foreign countries. No allowance is made, for example, for state or country price levels or for barriers to trade in the destination countries.[18] Hence, the model is reestimated using state-specific dummy variables, using destination-country-specific dummy variables, and using both sets of fixed effects.[19] This is done for the model that incorporates the corruption, language, and institutional similarity measures; the results are given in columns 3, 4, and 5.

[16] For example, see Hutchinson (2002).

[17] In our sample the English-speaking countries are Canada, Belize, Jamaica, Trinidad, Grenada, Guyana, the United Kingdom, Ireland, Singapore, Fiji, the Philippines, Hong Kong, Australia, New Zealand, Western Samoa, Cameroon, Ghana, Nigeria, Uganda, Kenya, South Africa, Zambia, and Zimbabwe. The Spanish-speaking countries are Mexico, Guatemala, El Salvador, Honduras, Nicaragua, Costa Rica, Panama, the Dominican Republic, Colombia, Venezuela, Ecuador, Peru, Bolivia, Chile, Paraguay, Uruguay, Argentina, and Spain. These groupings are based on the language in question being reported by the CIA Factbook as either the principal or, in the case of Hong Kong, second language of the particular country. The institutionally similar countries are Australia, New Zealand, Canada, Ireland, and the United Kingdom. The estimated results are insensitive to any reasonable alternative grouping.

[18] See Anderson and van Wincoop (2003) for an example in which calculation of an index of trade resistance is performed. See also Feenstra (2004, p. 161 ff.) for a discussion of the use of fixed effects in situations of this sort.

[19] Because our observations are all for a single time period, we cannot follow the stronger advice of Cheng and Wall (2002) to use trading-pair-specific bilateral fixed effects.

Column 3 reports the results of estimating the model with state fixed effects. The results are very similar to what is reported in column 2. The variable *MS* has a strong protrade effect, and the roles of *CORRUPTION* and of *ENGLISH* and *SPANISH* are confirmed. State-specific heterogeneity exists, but it does not alter the links between migrant stock and state exports. The basic proposition and its first two corollaries are supported.[20]

Next the model is modified to include destination-country fixed effects. Collinearity precludes the use of variables representing characteristics of the destination countries; hence, the country-specific fixed-effects model is the analog of either of the models in columns 1 and 2. Column 4 reports this estimation. Inclusion of the country-specific dummies has little effect on the estimated protrade elasticity of immigrants on exports, which remains a statistically and economically meaningful 0.24. The remaining control variables, other than *DISTANCE,* whose coefficient more than doubles in absolute value, are little changed.

Inclusion of both the state-specific and country-specific dummy variables yields the results reported in column 5. The only variables whose coefficients can be estimated here are *MS* and *DISTANCE*. The estimate on *MS* is 0.39, of the same order of magnitude obtained in the other specifications, and it remains statistically significant. Again, the basic proposition is supported.

A final exercise is to estimate the size of the immigrant stock that yields the peak protrade effect. Gould (1994) entered *MS* into his model so as to force a diminishing marginal effect. For overall exports from the United States as a whole, he found that 90% of the immigrant effect was exhausted at approximately 12,000 immigrants *in the entire United States* (p. 310). In our study the logarithm of *MS* was replaced by a quartic polynomial in *MS*. Doing this alters none of the conclusions reached earlier, and this approach indicates that the effects of immigrants attains a maximum for 6,200 to 7,400 immigrants *per state*.

V. Conclusions

Our results confirm the basic proposition that immigrants have a protrade effect on exports from the American states. We also go beyond the basic proposition to test three corollaries and find that a higher level of corruption in the destination country strengthens the protrade role of immigrants, although it otherwise leads overall to reduced trade. We also find that language similarity between the United States and the export destination country reduces the protrade effect of migrants, although language similarity otherwise promotes trade. Institutional similarity is not found to be important for the protrade effect of immigrants in the American states. These results are undiminished by inclusion of state or destination fixed effects.

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[20] In unreported estimations the roles of the Mexican and Canadian borders were considered, with little effect on the results.

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APPENDIX

Data

$EXPORTS_{ij}$: The average current dollar value of exports of manufactures from state i to country j for the years 1990, 1991, and 1992. The data are from the proprietary WISER Origin of Movement series.

MS_{ij}: Migrant stock. The number of people born in country j residing in state i as recorded in the 1990 Census. From Lapham (n.d.).

GSP_i: Gross state product of state i in 1990, from the U.S. Department of Commerce.

$PCGSP_i$: Per capita GSP of state i. The population data are from the 1990 U.S. Census.

GDP_j: Gross domestic product of destination country j in 1990, from the International Monetary Fund, *International Financial Statistics*.

$PCGDP_j$: Real per capita income of export destination country j in 1990, from Penn World Table, Series GRIP: given by real GDP per capita in 1985 international prices.

$CORRUPTION_j$: Proprietary data from the International Country Risk Guide, PRS Group. It is an index of corruption in government. The data are for 1991. Our variable is the negative of the deviation of the PRS index from its own mean.

$ENGLISH_j$ $(SPANISH_j)$: English (Spanish) language. Equals 1 if the country of export destination is significantly English-(Spanish-) speaking; otherwise it equals 0. The on-line CIA Factbook is the source of these data. Lists of countries so identified are given in the text.

INSTITUTIONAL SIMILARITY$_j$: Equals 1 for Ireland, Canada, Australia, New Zealand, and the United Kingdom; otherwise it equals 0.

$DISTANCE_{ij}$: The great circle distance in miles, from Bali Online (www.indo.com/distance). It is measured between the principal, or central, cities of each U.S. state and each of the 87 partner countries.

THE EMPIRICAL ASSESSMENT OF TECHNOLOGY DIFFERENCES: COMPARING THE COMPARABLE

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Abstract—Since the first statement of Hicks's induced innovation hypothesis in 1932, a large number of theoretical and empirical studies have analyzed the issue of price-induced technological change—many of them on the basis of substitution elasticities. This note compares technologies across space and time on the basis of factual and counterfactual substitution elasticities and argues that differences in estimated substitution elasticities should be decomposed into two counterfactual components. The first component is designed to indicate how the ease of substitution is altered by varied economic circumstances; the second addresses the question of how technologies would compare under genuinely comparable situations. This argument is illustrated by the example of energy-price elasticities of capital before and after the oil crisis of the early 1970s.

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I. Introduction

As more and more evidence points to the existence of potentially substantial climate change, price-induced technological change has increasingly attracted attention from environmental economists due to its possibly ameliorating role regarding environmental problems. In particular, Popp (2002, p. 160) recently found that energy prices have strongly significant positive effects on energy-saving innovation. Within a product-characteristics framework, Newell, Jaffe, and Stavins (1999) also provide empirical evidence that energy

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