THE ECONOMIC VALUE OF CULTURAL DIVERSITY: EVIDENCE FROM US CITIES

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Abstract

We use data on wages and rents in different U.S. cities to assess the amenity effects on production and consumption of cultural diversity as measured by diversity of countries of birth of city residents. We show that US-born citizens living in metropolitan areas where the share of foreign-born increased between 1970 and 1990 have experienced a significant average increase in their wage and in the rental price of their housing. Such finding is economically significant and robust to omitted variable bias and endogeneity bias. We then present a model in which cultural diversity may have both production and consumption amenity or disamenity effects. As people and firms are mobile across cities in the long run, the model implies that the joint results from the wage and rent regressions are consistent with a dominant production amenity effect of cultural diversity.

JEL Classification: O4, R0, F1.

Keywords: cultural diversity, productivity, local amenities, urban economics.

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1 Introduction

In recent years there has been a resurgence of international migration directed to industrialized countries. As a consequence, the share of foreign-born residents has increased dramatically in the population of traditionally 'receiving countries', such as the United States, as well as of several European countries (most notably, France, Germany, and the UK; more recently, Italy, Spain, and Austria). Rising immigrant pressures in industrialized countries have generated an intense policy debate on the opportunity of imposing additional restrictions on legal and illegal migration flows.

The debate has been accompanied by a large empirical literature on the consequences of migration (see, e.g. Borjas, 1994, 1995; Boeri, Hanson and McCormick, 2002, Card 1990, 2001). Such literature has mostly focused on the short-run distributional consequences of migration in terms of lower wages and higher unemployment for unskilled natives, and on the rising costs of social security resulting from the inflow of relatively unskilled labor. A similar emphasis also characterizes the discussion on the long-run consequences of immigration that has been mainly framed within the neoclassical growth model (see, e.g., Dolado et al., 1994; Barro and Sala-i-Martin, 1995; Canova and Ravn, 2000). From such perspective, international immigration is assimilated to an increase in the rate of growth of the unskilled labour force resulting in a dilution of physical and human capital in the receiving countries. Migration has been studied as a mechanism that fosters convergence in income per capita and wages between capital-abundant receiving countries and capital-scarce sending ones.

Our work takes a different angle in looking at this issue. Rather than studying the short-run effects of new immigration on the receiving country in a classic model of skill supply and demand, we consider a multi-city model of production and consumption and we ask what is the value of the cultural diversity that foreign-born bring to each city. If cultural diversity is a city characteristic (certainly endogenous) we can learn about its value from the long-run equilibrium distribution of wages and prices across cities.

Diversity over several dimensions has been praised by economists as valuable in consumption and production. Jacobs (1969) attributes the success of cities to their industrial diversity. Glaeser et al. (2001) identify in the diversity of available consumption goods one of the attractive features of cities. More generally, Fujita et al (1999) use 'love of variety' in preferences and technology as the building block of their theory of spatial development.

We believe that cultural diversity may very well be an important aspect of diversity with consequences on production as well as consumption. The aim of this paper is to quantify the value of cultural diversity, as measured by the presence of foreign-born in a city, to US-born people. "It's hard to put a number on buzz but there must be some value" (Richard Freeman, cited by The Economist, 2002). Who can deny that Italian restaurants, French beauty shops, German breweries, Belgian chocolate stores, Russian ballets, Indian tea houses and Thai massages constitute valuable consumption amenities inaccessible to Americans were not for their foreign-born residents? Similarly the skills and abilities of foreign-born workers and thinkers may complement those of native workers and thus boost problem solving and efficiency on the workplace.[1] Cultural diversity would act as a production amenity in this case. On the other hand, natives may not like to live in a multicultural environment in so far as this may endanger their own cultural values or intercultural frictions may reduce their productivity. Cultural diversity would, then, act as a consumption or production disamenity respectively.

We focus on cities in the US as a natural laboratory, the reason being that cultural diversity has long been one of the hallmarks of US society. For this reason, our analysis on US cities serves as a benchmark for studies on other developed countries in Europe and Asia that are becoming increasingly diverse due to recent inflows of foreign workers. As US-born people are highly mobile across US cities, following Roback (1982) we develop a model of a multicultural system of open cities that allows us to use the observed variations of wages and rents of US-born workers to identify the nature of the externalities associated with cultural diversity. Our main finding is that, on average, *US-born citizens attribute a dominant production amenity value to cultural diversity.* We believe that this result is interesting, robust and new in the literature.

The rest of the paper is organized as follows. Section 2 reviews the literature on the economic consequences of cultural diversity. Section 3 introduces our dataset of 160 US metropolitan areas during the period 1970-1990 and surveys the main stylized facts: cultural diversity in a city is significantly positively correlated with the average wage and rent of US-born citizens in that city. Section 4 develops the theoretical model that is used to design our estimation strategy in terms of joint wage and rent equations. Section 5 runs the regressions and checks the results for robustness and endogeneity. Section 6 discusses the results and concludes.

2 Literature on Diversity

Cultural diversity and its effects, often defined in specific ways, have attracted the attention of many applied economists for a long time. The applied 'labor' literature has analyzed ethnic diversity and ethnic 'segregation' in the U.S. as well as its impact on economic discrimination and the achievements of minorities. The focus of attention has often been the black-white gap. Few examples among many contributions are Card and Krueger (1992), (1993), Cutler and Glaeser (1997), Arrow (1998), Eckstein and Wolpin (1999), Mason (2000). While the black-white issue can be reduced to different 'countries of origin' going far back in the past, this paper does not focus on this aspect of cultural diversity. We control for black-white composition issues but we never focus on them.

[1] The anedoctical evidence of the contribution of foreign born to 'big thinking' in the US is quite rich. One striking example is the following. In the last ten years, out of the 47 US-based Nobel laureates in Chemistry, Physics and Medicine, 25 per cent (14 laureates) were not US-born. During the same time period the share of foreign-born in the general population was on average only 13 per cent. From our perspective, such example is interesting because research in hard sciences is typically based on large team work.

Much more closely related to our analysis is the literature on the impact of immigration on the US labor market. Several contributions by George Borjas (1994), (1995), and (1999) focus on the issue of new immigrants into the US and their effect on native workers. Similarly important contributions by David Card (notably, Card, 1990; Card and Di Nardo, 2000; Card, 2001) analyze the reactions of domestic workers and their wages to inflows of new immigrants. These contributions do not seem to achieve a consensus view either on the effect of new immigrants on wages of low skilled domestic workers (which seems, however, small) or on the effect of new immigrants on the migration behavior of domestic workers. More recently, quite convincing evidence of a positive effect of immigrant inflows on rents in cities has been provided by Saiz (2003a,b). All these studies share some common features especially in terms of their methodological approach. They all focus on the impact of new immigrants on wages (rents) and domestic migration in the short run (within years) and use a classic frame of labor demand-supply to analyze the effects. They assume that immigrant and domestic workers, within a skill group, are homogeneous so that immigration is a shift in labor supply, which affects local wages (rents) more or less depending on the mobility of domestic workers. Our approach takes a rather different stand. We consider that being 'foreign-born' is a feature that permanently differentiates individuals (either new or old immigrants) in terms of their non-market attributes and such feature may have consumption and production amenity (or disamenity) value for US-born workers. Moreover, we consider long-run variations of wages and rents relying on the assumption of perfect mobility of US-born workers and firms across cities in the long run.

Fewer contributions have focused on other aspects of diversity (cultural and linguistic) or looked at its relationships with productivity and welfare of US-born people. Most of the studies focus on the downside of diversity in terms of its static costs associated with lack of communication and transaction barriers. For example, Lazear (1995) assumes that a common culture and a common language facilitate exchange and trade between individuals. He argues that minorities have incentives to become assimilated and to learn the language of the majority in order to participate into a larger pool of potential trading partners. In his model, as individuals do not properly internalize the social value of assimilation, multiculturalism is bad. Alesina, Baqir and Easterly (1999) look at the relation between the heterogeneity of preferences and the provision of public goods in US cities. They show that the share of spending on productive public goods is inversely related to the ethnic fragmentation of cities even after controlling for other socioeconomic and demographic determinants. Here again cultural diversity is bad.

Interestingly, and related to our work, several researchers in social sciences have related diversity with urban agglomerations. The functioning and thriving of urban clusters seem to rely on the effective interaction of many units which are 'diverse' in many respects. A first example is given by urban studies. Jacobs (1969) sees economic diversity as the key factor of a city's success. Sassen (1994) studies 'global cities' - such as London, Paris, New York, and Tokyo - and their strategic role in the development of activities that are central to world economic growth and innovation, such as finance and specialized services. A key feature of these cities is the cultural diversity of their populations. Similarly, Bairoch (1998) sees cities and their diversity as the engine of economic growth. Such diversity, however, has been mainly investigated in terms of diversified provision of consumers' goods and services as well as productive inputs (see, e.g., Quigley, 1998; Glaeser et al., 2001). In his work at the interface between sociology and economics, Richard Florida (2003) argues that 'diverse' and tolerant cities, populated by artists, bohemians, and other creative people are also the most innovative cities in terms of high tech sectors. Our analysis of the role of cultural diversity is an extension of these lines of research.

Another literature is also potentially relevant to our work in that it motivates the positive 'production value' of diversity. It consists of studies on the organization and the management of teams. A standard assumption is that diversity leads to more innovation and creativity because diversity implies different ways of framing problems, a richer set of alternative solutions, and therefore higher quality decisions. Lazear (1999) provides an attempt to model team interactions. He defines the 'global firm' as a team whose members come from different cultures or countries. Combining workers who have different cultures, legal systems, and languages imposes costs on the firm that would not be present if all the workers were similar. However, complementarity between workers, in terms of disjoint and relevant skills, offsets the costs of cross-cultural interaction.[2] Here, again, multiculturalism is good.

Finally, there is a strand of studies in political economics that looks at the historical effects of cultural and ethnic diversity on the formation and the behavior of institutions. Across countries, the extent of government corruption, bureaucratic red tape, and black market activities as well as the protection of property rights seem to be all affected by the degree of ethnical fragmentation. The traditional wisdom (confirmed by Easterly and Levine, 1997) used to be that more fragmented (i.e. diverse) societies promote more conflict and predatory behavior, and generate less growth. However, recent studies have questioned that logic by showing that higher ethnic diversity is not harmful to economic development (see, e.g., Liam and Oneal, 1997). Collier (2001) actually finds that, as long as their institutions are democratic, fractionalized societies have better economic performance in their private sector than more homogenous ones. In our work we take institutions as given and equal across US cities and we only look at the effect of diversity on production and consumption within such institutional framework. It is interesting to notice, however, that also from a historical perspective the issue of how diversity affects productivity and development is still somewhat controversial.

[2] Fujita and Berliant (2003) model 'assimilation'as a result of team work: the very process of cooperative knowledge creation reduces the heterogeneity of team members through the accumulation of knowledge in common. Under this respect, a perpetual reallocation of members across different teams may be necessary to keep creativity alive.

3 Cultural Diversity, Wages and Rents

This paper takes a very US-based approach to the issue of cultural diversity. The question we are interested in is: What is there in cultural diversity for the US-born people ? Do they benefit at all from the presence of foreign-born? Do they value it? If they do, how do we measure such benefits?

Our analysis extracts the answers to those questions from the equilibrium outcome deriving from the implicit evaluation of diversity that the US-born make by 'voting with their feet'. The underlying assumption is that US-born workers and US firms are very mobile across cities in the long run. This assumption is motivated by extensive empirical evidence that shows very large gross migration flows across states and cities. For instance, using census data, we calculate that 36% of the population moved from one state to another between 1985 and 1990. As people respond to changes in the local working and living environment of cities, the wage and rent variations that we observe in the long run should reflect a spatial equilibrium: workers and firms are indifferent among alternative locations because they have eliminated any systematic difference in indirect utility and profits through migration.[3] While postponing the formalization of these ideas to Section 4, here we introduce our measure of cultural diversity and present some suggestive stylized facts about its relationship with average wages and rents in US cities.

3.1 Data and Diversity Index

Data at the Metropolitan Statistical Area (MSA) level for the United States are available from different sources. We use mostly the Census Public Use Microdata Sample (PUMS) for year 1971 and 1991 in order to calculate wages and rents for specific groups of citizens in each MSA. We use the 1/100 sample from the 15% PUMS of 1970 and the 5% PUMS for 1990. We also use data from the 'County and City Data Book' from several years in order to obtain some aggregate variables such as employment, income, population, spending for local public goods. We consider 160 Standard MSAs (SMSAs) that are identified in each of the census years considered. We have around 1,200,000 individual observations for 1990, and 500,000 for 1970. We use them to construct aggregate variables and indices at the SMSA level. The reason for focusing on SMSAs is twofold. First, SMSAs constitute closely connected economic units within which interactions are intense. Thus, they seem to fit our theoretical model in which commuting takes place within cities but not between cities. Second, they exhibit a higher degree of diversity than the rest of the country as new immigrants and their offsprings traditionally settle down in larger cities.

We measure the average wage of native workers in an MSA using the yearly wage of white US-born male residents between 40 and 50 years of age. The average yearly wage constructed using this procedure for city c in year t, call it $\overline{w}_{ct}$ with $c = 1, ..., 160$, is neither affected by composition effects nor distorted by

[3] We are grateful to Ed Glaeser for drawing our attention to the potential dividends of this approach.

potential discrimination factors (across genders or ethnicity) and it is therefore a good proxy of the average wage of US-born workers in the city. In particular, the construction of $\overline{w}_{ct}$ is not affected by the degree of diversity of a city. The correlation between diversity and $\overline{w}_{ct}$ comes only from the equilibrium effect of diversity on labor demand and labor supply.

As measure of the average land rent in a MSA we use the average monthly rent paid per room in the house (i.e., the monthly rent divided by the number of rooms) by white US-born people in working age (16-65). We call $\overline{r}_{ct}$ such measure for city c in year t.

Turning to our key explanatory variable, our measure of cultural diversity considers the country of origin of people as defining their cultural identity. Cultural diversity is certainly a multidimensional concept and could stem from different ethnicity, religion, national origin or other characteristics. Here, however, we focus on differences in country of birth as such diversity is likely to increase as a result of migration and it is highly correlated with linguistic and national identity. foreign-born have always been an important share of the US population and their proportion has been growing in the past decades. In 1970 they were 8 percent of the total working age population. In 1990 they reached 12 percent and they kept on growing afterwards.

To keep our dataset comparable with existing cross-country studies, we use a rather standard measure of diversity, namely, the so called 'index of fractionalization' (henceforth, simply 'diversity index'). Such index has been popularized in cross-country studies by Mauro (1995) and largely used thereafter. It is nothing but the Simpson index used to measure biodiversity and embeds the probability that two randomly selected individuals in a community belong to different groups. It accounts for the two main dimensions of diversity, i.e., 'richness' (number of groups) and 'evenness' (balanced distribution of individuals across groups).[4] The index is calculated as 1 minus the Herfindal index of concentration across groups. Specifically, in the case of the variable CoB (country of birth) the corresponding index is defined as:

$$div(CoB)_{ct} = 1 - \sum_{i=1}^{M} (CoB_i^c)_t^2 \tag{1}$$

where $(CoB_i^c)_t$ is the share of people born in country i among the residents of city c in year t. This index reaches its maximum value 1 when each individual is in a different group, i.e. there are no individuals born in the same country, and its minimum value 0 when all individuals belong to the same group, i.e. all individuals were born in the same country. However, in measuring diversity there is something specific to our data set. First, in each city the largest group, by far, is always represented by the US-born. Second, most of the variation across cities in $div(CoB)_{ct}$ depends on the variation of the share of foreign-born $(Foreign_c) = \sum_{i \neq US}^{M} (CoB_i^c)_t$ rather than on the variation in the countries of

[4] Despite differences that may seem notable at first sight, most statistical measures of diversity are either formally equivalent or at least highly correlated when run on the same data set. See Maignan et al (2003) for details.

origin. On both counts, an alternative, and sometimes preferable, measure of diversity could simply be the share of foreign-born. We will present results using both measures.

The 1970 and 1990 PUMS data report the country of birth of each individual. We consider as separate groups each country of origin of migrants contributing at least 0.5 percent of the total foreign-born living and working in the US. The other countries of origin are gathered in a residual group. Such choice implies that we consider 35 countries of origin in 1970 as well as in 1990. Such choice covers about 92 percent of all foreign-born immigrants while the remaining 8 percent are merged into one group. The complete list of countries for each census year is reported in the data appendix and the largest 15 of these groups are reported in Table 1. As the table shows, between 1970 and 1990, the origin of migrants has become increasingly polarized towards Mexican immigrants, but the share of foreign-born has increased so that, overall, the diversity index has increased. As to the main sources of immigrants, we also notice the well known shift from European countries to Asian and Latin American countries.

Table 1
Foreign Born living in 161 U.S. MSA's,
15 Largest Groups 1970-1990

Country of Origin	Percentage of total Foreign Born 1970	Country of Origin	Percentage of total Foreign Born 1990
Canada	9.0%	**Mexico**	20.0%
Italy	8.1%	**Philippines**	6.0%
Germany	7.8%	**Cuba**	4.2%
Mexico	7.3%	**Germany**	3.2%
Syria	7.0%	**Canada**	3.2%
Cuba	5.1%	**China**	2.8%
Poland	4.5%	**India**	2.8%
UK	4.4%	**Viet-Nam**	2.7%
Philippine	2.3%	**El Salvador**	2.6%
USSR	2.3%	**Italy**	2.4%
Ireland	2.3%	**Korea**	2.2%
China	2.3%	**UK**	2.2%
Yugoslavia	1.7%	**Japan**	1.8%
Greece	1.6%	**Jamaica**	1.7%
Hungary	1.6%	**Colombia**	1.6%
Foreign Born as % of working age total population , 1970	8.0%	**Foreign Born as % of working age total population , 1990**	11.9%

Source: Authors' Elaborations on 1970 and 1990 PUMS Census Data.

3.2 Diversity Across U.S. Cities

In order to convince the reader that US cities are a very differentiated universe in terms of diversity and that there is enough variation across them to be able to learn something precise from their analysis, Table 2 shows the percentage of foreign-born and the Diversity Index for a group of important Metropolitan areas.

To put into context the extent of diversity in US cities, their diversity can be compared with the cross-country values of the index of linguistic fractionalization reported by the Atlas Narodov Mira and published in Taylor and Hudson (1972) for year 1960. Such values have been largely used in the growth literature (see, e.g., Easterly and Levine, 1997, and Collier, 2001). As foreign-born immigrants normally use their country's mother tongue at home and in turn this signals their country's cultural identity, our diversity index captures cultural and linguistic fragmentation just as that index does at the country level. The comparison yields intriguing results. Diversified cities, such as New York or Los Angeles, have diversity indices in the range from 0.5 to 0.6, which are comparable to the values calculated for countries such as Rhodesia (0.54), which is often disrupted by ethnic wars, or Pakistan (0.62), which also features a problematic mix of conflicting cultures. Afghanistan, a well known quagmire of different cultural identities, reaches a value of 0.66 that is only slightly higher. More homogenous cities, such as Cincinnati and Pittsburgh, exhibit a degree of fractionalization equal to 0.05, which is the same as that of very homogenous European countries, such as Norway or Denmark in the sixties. Between these two extremes *US cities span a range of diversity that is about two thirds of the range spanned by countries in the world.* Table 2 also shows that, even though people born in Mexico constitute an important group in many cities, the variety of origin of the foreign-born migrants across US cities is still remarkable.

Finally, from Table 2 we also get the impression of a very high positive correlation between the share of foreign-born people in a city and its diversity index. This confirms what was anticipated above: the presence of a large share of foreign-born, more than their group composition, is the largest source of diversity in US cities. Over the whole sample of 160 MSAs, the correlation coefficient between the two measures is 0.86 for 1990 and 0.87 for 1990. Similarly, for the 1970-90 period the correlation coefficient between the increase in the share of foreign-born and the increase in the diversity index is 0.84. Differently, the correlation between the increase in the diversity index calculated for the whole population and the diversity index calculated only within the group of foreign-born is a mere 0.08.

Table 2
Diversity in some U.S. MSA's, 1990

city	Share of Foreign Born	Country of Origin of the Five Largest foreign Groups	Diversity Index
Atlanta, GA	**5.8%**	**Germany, Mexico, India, England, Korea**	**0.11**
Chicago, IL	**15.2%**	**Mexico, Poland, Philippines, India, Germany**	**0.28**
Cincinnati, OH-KY-IN	**2.3%**	**Germany, England, India, Canada, Viet-Nam**	**0.057**
Dallas, TX	**10.6%**	**Mexico, Salvador, Viet-Nam, India, Germany**	**0.20**
El Paso, TX	**29%**	**Mexico, Japan, Korea, Canada, Panama**	**0.43**
Indianapolis, IN	**2.3%**	**Germany, England, Korea, Canada, Philippines**	**0.046**
Las Vegas, NE	**12%**	**Mexico, Philippines, Germany, Canada, Cuba**	**0.23**
Los Angeles, CA	**37%**	**Mexico, Salvador, Philippines, Guatemala, Korea**	**0.58**
New York, NY	**31%**	**Dominican Republic, China, Jamaica, Italy, Colombia**	**0.51**
Oklahoma City, OK	**4.1%**	**Mexico, Viet-Nam, Germany, England, Japan**	**0.08**
Philadelphia, PA-NJ	**5%**	**Germany, India, Italy, England, Philippines**	**0.10**
Pittsburgh, PA	**2.3%**	**Italy, Germany, India, England, Canada**	**0.04**
Sacramento, CA	**10.6%**	**Mexico, Philippines, Germany, China, Canada**	**0.19**
San Francisco, CA	**30.3%**	**Philippines, China, Mexico, Salvador, Hong Kong**	**0.50**
Washington, DC-MD-VA-WV	**14.8%**	**Salvador, Germany, India, Korea, Viet-Nam**	**0.27**

Source: Authors' Elaborations on 1970 and 1990 PUMS Census Data.

3.3 Stylized Facts

Previewing the final results of our analysis, the key empirical finding is readily stated: keeping every other city characteristic equal, on average *US-born workers living in cities with richer cultural diversity are paid higher wages and pay higher rents than those living in cities with poorer cultural diversity.* Our main effort in section 5 will be to show that not only is this correlation not driven by any other (omitted) variable, but it is the result of causation running from diversity to wages and rents. To support such effort, section 4 will develop a theoretical model arguing that, when firms and workers are freely mobile across cities, the above finding can be explained in equilibrium only if diversity has a *dominant production amenity effect.* As a natural first step in that direction, the present subsection reports the raw correlations between diversity on the one hand and wages as well as rents on the other.

While there is a strong positive correlation in the cross section, it is more effective to report the correlation across the 160 cities between the change, from 1970 to 1990, of the share of foreign-born, $\Delta(Foreign_c)$, and the percentage change in the wage of the US-born, $\Delta \ln(\,\overline{w}_c)$, or the percentage change in rents paid by the US-born, $\Delta \ln(\,\overline{r}_c)$. Any fixed characteristic of cities, such as their location, their fixed amenities (e.g., weather conditions), and their traditions, does not affect that correlation and this is why we prefer it as stylized fact. Figure 1 and 2 show the scatter-plots for these two partial correlation and report the OLS regression. Cities that have substantially increased their share of foreigners in the twenty years considered, such as Los Angeles, Miami and San Francisco, have experienced increases in wages larger than average; similarly (with the exception of Miami) Los Angeles, San Francisco, San Jose, and Jersey City, which experienced a large increase in foreign-born, also experienced large increases in rents. The OLS coefficient implies that a 10 percent increase of the share of foreign-born (the difference between, say, Oklahoma City and Chicago in 1990) is associated with 7 percent higher wages and 11 percent higher rents for same sized houses. Similar estimates would be obtained using the 1990 cross section (rather than changes): a difference of 10 percent in foreign-born people would be associated with differences of 9 percent in wages and 15 percent in rents. The raw positive correlation is very significant. Let us anticipate that, even when controlling for increases in average schooling and for the change in city-employment, the effect on wages still stands at 5.6 percent (standard error 1.6 percent). Analogously, when controlling for the changes in city population density and in per-capita income, the effect on rents still stands at 5.3 percent (standard error 2 percent).

Figure 1 - Wages and Diversity

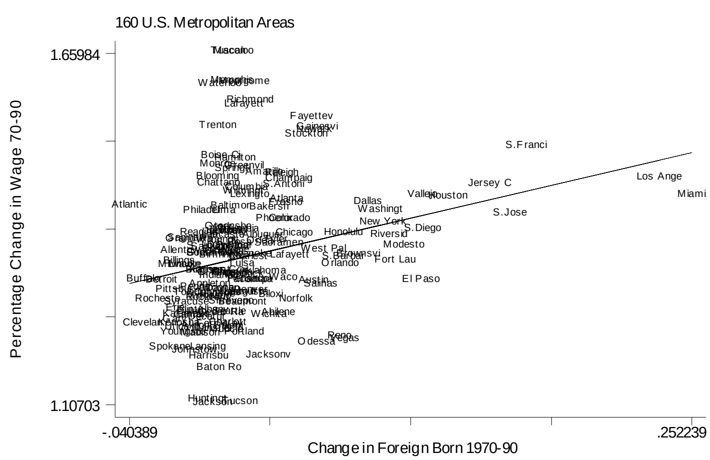
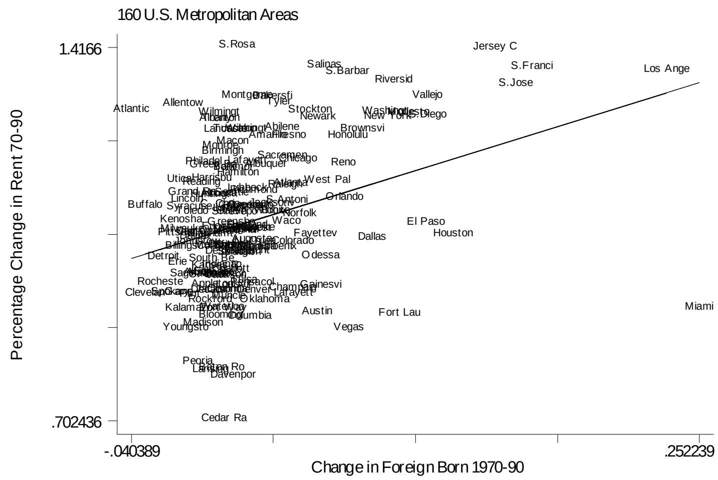


Figure 2 - Rents and Diversity



4 A Multicultural City System

To structure our empirical investigation, we develop a stylized model of an open system of cities in which 'diversity' affects both the productivity of firms and the satisfaction of consumers through a localized external effect. Both the model and the identification procedure of the impact of diversity on city dwellers build on Roback (1982).

4.1 The Model

We consider an open system of a large number N of non-overlapping cities, indexed by $c = 1, ..., N$. There are two factors of production, labor and land. Labor is perfectly mobile between and within cities. We assume that intercity commuting costs are prohibitive so that for any worker the cities of work and residence coincide. We also ignore intra-city commuting costs, which allows us to focus on the intercity allocations of workers.

The overall amount of labor available in the economy is equal to L. It is inelastically supplied by urban residents and, without loss of generality, we choose units such that each resident supplies one unit of labor. Accordingly, we call L_c the number of workers employed and resident in city c. Workers are all identical in terms of attributes that are relevant for market interactions. However, they differ in terms of non-market attributes, which exogenously classifies them into M different groups ('cultural identities') indexed by $i = 1, ..., M$. Hence, calling L_i the overall number of workers belonging to group i, we have $\sum_{i=1}^{M} L_i = L$. In each city cultural diversity d_c, measured in terms of the number ('richness') and relative sizes L_{ic} ('evenness') of resident groups, enters both production and consumption as an externality that, in principle, can be positive ('amenity') or negative ('disamenity'). To establish the existence and the sign of such externality is the final aim of the paper. Differently from labor, land is fixed among cities. It is nonetheless mobile between uses within the same city. We call H_c the amount of land available in city c. As to land ownership, we assume that the land of a city is owned by locally resident landlords.[5]

To summarize, while the intercity allocation of land is exogenously given, the intercity allocation of labor will be endogenously determined in equilibrium. Accordingly, while the city system as a whole is characterized by an exogenous degree of cultural diversity, within city diversity is endogenously determined by the entry decisions of firms and the migration decision of workers.

Preferences are defined over the consumption of land H and a homogeneous good Y that is freely traded among cities. Specifically, the utility of a typical worker of group i in city c is given by:

$$U_{ic} = A_U(d_c)\, H_{ic}^{1-\mu} Y_{ic}^{\mu} \tag{2}$$

with $0 < \mu < 1$. In (2) H_{ic} and Y_{ic} are land and good consumption respectively while $A_u(d_c)$ captures the consumption externality associated with local diversity d_c. If the first derivative $A_u'(d_c)$ is positive, diversity is a consumption amenity; if negative it is a consumption disamenity.

We assume that workers move to the city that offers them the highest indirect utility. Given (2), utility maximization yields:

$$r_c H_{ic} = (1-\mu) E_{ic},\ p_c Y_{ic} = \mu E_{ic} \tag{3}$$

which implies that the indirect utility of the typical worker of group i in city c is:

$$V_{ic} = (1-\mu)^{1-\mu} \mu^{\mu}\, A_u(d_c)\, \frac{E_{ic}}{r_c^{1-\mu} p_c^{\mu}} \tag{4}$$

where E_{ic} is her expenditures while r_c and p_c are the local land rent and good price respectively.

As to production, good Y is supplied by perfectly competitive firms using both land and labor as inputs. The typical firm in city c produces according to the following technology:

$$Y_{jc} = A_Y(d_c)\, H_{jc}^{1-\alpha} L_{jc}^{\alpha} \tag{5}$$

with $0 < \alpha < 1$. In (5) H_{jc} and L_{jc} are land and labor inputs respectively while $A_Y(d_c)$ captures the production externality associated with local diversity d_c. If $A_Y'(d_c)$ is positive, diversity is a production amenity; if negative it is a production disamenity.

[5] This assumption is made only for analytical convenience. What is crucial for what follows is that the rental income of workers, if any, is independent of locations and, thus, it does not affect the migration choice. The alternative assumptions of absentee landlords or balanced ownership of land across all cities would also serve that purpose.

Given (5) and perfect competition, profit maximization yields:

$$r_c H_{jc} = (1-\alpha) p_c Y_{jc}, \ w_c L_{jc} = \alpha p_c Y_{jc} \tag{6}$$

which implies marginal cost pricing:

$$p_c = \frac{r_c^{1-\alpha} w_c^{\alpha}}{(1-\alpha)^{1-\alpha} \alpha^{\alpha} A_Y(d_c)} \tag{7}$$

so that firms make no profits in equilibrium. Given our assumption on land ownership, this implies that aggregate expenditures in the city equal local factor incomes and that workers' expenditures consist of wages only: $E_{ic} = w_c$. As good Y is freely traded, its price is the same everywhere. We choose the good as numeraire, which allow us to write $p_c = 1$.[6]

In a spatial equilibrium there exists a set of prices $(w_c, \ r_c, \ c = 1, ..., N)$ such that in all cities workers and landlords maximize their utilities given their budget constraints, firms maximize profits given their technological constraints, factor and product markets clear. Moreover, no firm has an incentive to exit or enter. This is granted by (7) that, given our choice of numeraire, can be rewritten as:

$$r_c^{1-\alpha} w_c^{\alpha} = (1-\alpha)^{1-\alpha} \alpha^{\alpha} A_Y(d_c) \tag{8}$$

We will refer to (8) as the 'free entry conditions'. Finally, in a spatial equilibrium no worker has an incentive to migrate. For an interior equilibrium (i.e., $L_c > 0$ $\forall c = 1, ..., N$) that is the case when workers are indifferent between alternative cities:

$$V_{ic} = V_{ik}, \ \forall c, k = 0, ..., N \tag{9}$$

We will refer to (9) as the 'free migration conditions'.

To conclude the solution of the model we have to determine the spatial allocation of workers L_{ic}. This is achieved by evaluating the implications of market clearing for factor prices. Specifically, given $L_c = \sum_j L_{jc}$ and $Y_c = \sum_j Y_{jc}$, (6) imply $w_c L_c = \alpha p_c Y_c$. Given $H_c = \sum_j H_{jc} + \sum_i H_{ic}$, (6) and (3) imply $\mu r_c H_c = (1-\alpha\mu) p_c Y_c$. Together with $E_{ic} = w_c$ and $p_c = 1$, those results can be plugged into (4) to obtain:

$$V_{ic} = \mu \left(\frac{1-\mu}{1-\alpha\mu} \right)^{1-\mu} \left(\frac{H_c}{L_c} \right)^{1-\alpha\mu} A_U(d_c) \left[A_Y(d_c) \right]^{\mu} \tag{10}$$

Substituting (10) into (9) completes the system of equations that can be solved for the equilibrium spatial allocation of workers. In particular, such substitution gives $M(N-1)$ free migration conditions that, together with the M group-wise full-employment conditions $\sum_{c=1}^{N} L_{ic} = L_i$, assign L_{ic} mobile workers of each group $i = 1, .., M$ to each city $c = 1, ..., N$.

[6] Anticipating the empirical implementation of the model, by setting $p_c = 1$ for all cities we are requiring the law-of-one-price to hold for tradable goods and non-tradable goods prices to be reasonably proxied by land rents. This seems to be supported by the large positive correlation between local price indices and land rents at the SMSA level.

Due to constant returns to scale and fixed land, (10) shows that the indirect utility offered to a worker in each city is a decreasing function of the total number of local workers. This ensures the uniqueness of the spatial equilibrium in terms of city sizes L_c's. Moreover, (10) also shows that the local indirect utility tends to infinity as all workers abandon a certain city, which ensures that the unique equilibrium has indeed a positive number of workers in every city ('no ghost town'). Finally, whether in equilibrium cities have a more or less diversified group composition (d_c high or low), depends on the combined consumption and production external effects of diversity $A_U(d_c)\,[A_Y(d_c)]^{\mu}$. If such combination generates a net amenity effect, cities will tend to be diversified; if it generates a net disamenity effect, they will tend to be homogeneous. More precisely, due to symmetry among groups, in the presence of a net amenity effect of diversity, cities will have a uniform distribution of workers across groups ('multicultural cities'); if a net disamenity effect arises, different groups will tend to concentrate in different cities ('unicultural cities').[7]

4.2 Identification: Wage and Rent Equations

To prepare the model for empirical investigation, it is useful to evaluate wages and land rents at the equilibrium allocation. This is achieved by solving together the logarithmic versions of the free entry condition (8) and the free mobility condition (9) that takes (4) into account. Specifically, call v the equilibrium value of indirect utility. Due to free mobility such value, call it v, is common among cities and, due to the large number of cities, it is unaffected by city-level idiosyncratic shocks. Then, solving (8) and (9) for factor prices gives the 'rent equation':

$$\ln r_c = \frac{\eta_Y + \alpha\eta_U}{1-\alpha\mu} + \frac{1}{1-\alpha\mu}\ln\left(A_Y(d_c)\,[A_U(d_c)]^{\alpha}\right) \tag{11}$$

and the 'wage equation':

$$\ln w_c = \frac{(1-\mu)\eta_Y - (1-\alpha)\eta_U}{1-\alpha\mu} + \frac{1}{1-\alpha\mu}\ln\left(\frac{[A_Y(d_c)]^{1-\mu}}{[A_U(d_c)]^{1-\alpha}}\right) \tag{12}$$

where $\eta_Y \equiv \ln(1-\alpha)^{1-\alpha}\alpha^{\alpha}$ and $\eta_U \equiv (1-\mu)^{1-\mu}\mu^{\mu}/v$.

Equations (11) and (12) constitute the theoretical foundations of our following regressions. They capture the equilibrium relation between diversity and factor prices. In the wake of Roback (1982), they have to be estimated together since they clearly show that any regression of one equation alone runs into a problem of lack of identification. To see this, consider (11) in isolation. A positive correlation between d_c and r_c is consistent with both a dominant consumption amenity effect of diversity ($A_U'(d_c) > 0$) and a dominant production

[7] In the case of net amenity, the multicultural equilibrium configuration is unique. In the case of net disamenity, the unicultural equilibrium configuration is unique if there are more cities than groups ($N \geq M$): any city hosts only one group of workers. On the contrary, if there are more groups than cities, $N < M$ multiple equilibria exist as some groups have to coexist within the same city.

amenity effect ($A'_Y(d_c) > 0$). Analogously, if one considers (12) in isolation, a positive correlation between d_c and w_c is consistent with both a dominant consumption disamenity effect ($A'_U(d_c) < 0$) and a dominant production amenity effect ($A'_Y(d_c) > 0$). Only the joint estimation of (11) and (12) allows one to establish which effect is indeed dominating. Specifically:

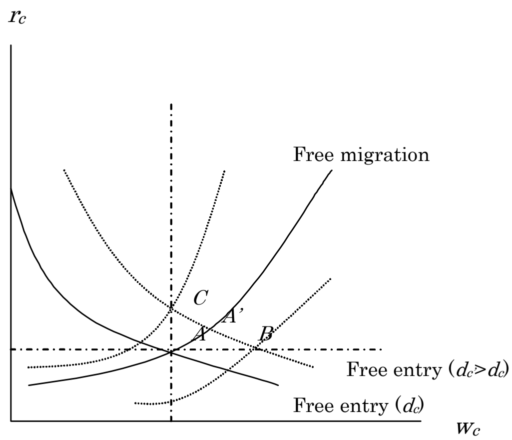
$$\begin{aligned}
\frac{\partial r_c}{\partial d_c} &> 0 \text{ and } \frac{\partial w_c}{\partial d_c} > 0 \text{ iff } \textit{dominant production amenity } \; (A'_Y(d_c) > 0) \quad (13)\\
\frac{\partial r_c}{\partial d_c} &> 0 \text{ and } \frac{\partial w_c}{\partial d_c} < 0 \text{ iff } \textit{dominant consumption amenity } \; (A'_U(d_c) > 0)\\
\frac{\partial r_c}{\partial d_c} &< 0 \text{ and } \frac{\partial w_c}{\partial d_c} < 0 \text{ iff } \textit{dominant production disamenity } \; (A'_Y(d_c) < 0)\\
\frac{\partial r_c}{\partial d_c} &< 0 \text{ and } \frac{\partial w_c}{\partial d_c} > 0 \text{ iff } \textit{dominant consumption disamenity } \; (A'_U(d_c) < 0)
\end{aligned}$$

Figure 3 provides a graphical intuition of the proposed identification. In the figure w_c and r_c are measured along the horizontal and vertical axes respectively. For given v and diversity d_c, the free entry condition (8) is met along the downward sloping curve, while the free migration condition (9) holds along the upward sloping curve. The equilibrium factor prices are found at the intersection of the two curves. Diversity d_c acts as a shift parameter on the two curves: any shock to diversity shifts both curves. An increase in d_c shifts (8) up (down) if diversity has a production amenity (disamenity) effect. It shifts (9) up (down) if diversity has a consumption amenity (disamenity) effect. Thus, by looking at the impact of a diversity shock on the equilibrium wage and rent, we are able to identify the dominant effect of diversity. For example, consider the initial equilibrium A and the new equilibrium A' that prevails after a shock to diversity. In A' both w_c and r_c have risen. Our identification argument states that both factor prices rise if and only if an upward shift of (8) dwarfs any shift of (9), i.e., the production amenity effect dominates.

With respect to Roback (1982), however, we face an additional problem. While her focus is on fixed amenities (e.g., "clean air", "lack of severe snow storms", Roback, 1982, p.1260-61), diversity in our model is endogenous since it is determined by the migration decisions of workers. This implies that, in order to test any causal relation from diversity to wages and rents, diversity has to be instrumented. We will take due account of this endogeneity problem in subsection 5.3.[8]

[8]See, e.g., Glaeser and Maré (2001) for a discussion of the many pitfalls in estimating wage equations when firms and workers are mobile.

Figure 3 – The Spatial Equilibrium



5 Wage and Rent Regressions

The theoretical model provides us with a consistent framework to structure our empirical analysis. In particular, in the wake of Roback (1982) it suggests how to use wage and rent regressions to identify the external effect of diversity on production and consumption.

5.1 Basic Specifications

Our units of observation are the 160 Metropolitan Statistical Areas (MSAs) listed in the Appendix. The years of observation are 1970 and 1990. As an empirical implementation of the wage equation (12), we run the following basic regression:

$$\ln(\overline{w}_{c,t}) = \beta_1 \left(\overline{s}_{c,t}\right) + \beta_2 \ln\left(Empl_{c,t}\right) + \beta_3 div(CoB)_{c,t} + e_c + e_t + e_{ct} \qquad (14)$$

The dependent variable $\overline{w}_{c,t}$ is the average wage in city c in year t. That is measured as the average wage of US-born white males between 40 and 50 years of age (see Section 3.1 for details). This allows us to avoid issues of age, gender, and race composition. The focal independent variable is $div(CoB)_{c,t}$, which is the diversity index defined in equation (1). The other independent variables are controls. Specifically, $\overline{s}_{c,t}$ measures the average years of schooling for the group of white US-born males aged from 40 to 50. $Empl_{c,t}$ is total non-farm employment in city c and year t. We control for unobserved factors that may vary across cites (and not over time) such as location, climate and traditions by including 160 city fixed effects e_c. We also control for common effects over time (such as the generalized increase in immigrants as well as in wages and rents) by including a year dummy e_t. Finally, e_{ct} is a zero-mean random error term independent from the other regressors.

Under this assumption, the coefficient β_3 captures the equilibrium effect on wages of a change in cultural diversity. However, as discussed in the subsection 4.2, the sign of β_3 cannot be directly interpreted as evidence of any amenity effect of diversity. The reason is that it may signal either a dominant positive external effect of diversity in production, which increases labor demand, or a dominant negative external effect of diversity in consumption, which decreases labor supply. Identification, thus, requires to estimate a parallel rent regression. Based on the rent equation (11), we run the basic regression:

$$\ln(\overline{r}_{c,t}) = \gamma_1 \ln(\overline{y})_{c,t} + \gamma_2 \ln\left(Pop_{c,t}\right) + \gamma_3 div(CoB)_{c,t} + \varepsilon_c + \varepsilon_t + \varepsilon_{ct} \qquad (15)$$

The dependent variable $\overline{r}_{c,t}$ is the average monthly rent per room paid by white US-born males in city c in year t. The focal independent variable is again the diversity index $div(CoB)_{c,t}$. The other independent variables are controls. Specifically, $(\overline{y})_{c,t}$ is the average yearly income of the group of white US-born males in city c in year t, while $Pop_{c,t}$ is the population density in the city. Again we control for city fixed effects ε_c, a year dummy ε_t, and we assume that ε_{ct} is a zero-mean random error uncorrelated with the regressors. The coefficient γ_3 captures the equilibrium effect of a change in cultural diversity on average city rents. Then, by crossing the information on the signs of β_3 and γ_3, we are able to use (13) to identify the net external effect of diversity: dominant production amenity if and only if $\beta_3 > 0$ and $\gamma_3 > 0$; dominant consumption amenity if and only if $\beta_3 < 0$ and $\gamma_3 > 0$; dominant production disamenity if and only if $\beta_3 < 0$ and $\gamma_3 < 0$; dominant consumption disamenity if and only if $\beta_3 > 0$ and $\gamma_3 < 0$.

The results of the regressions (14) and (15) are reported in Table 3 and 4 respectively. These results are obtained by OLS estimation with city and time fixed effects while correcting the standard errors to be heteroskedasticity robust. In Table 3, Specification 1 is exactly the one described in (14). Returns to one year of schooling are estimated around 10 percent and the change in employment is not significantly correlated with wages. The diversity index has a positive and very significant effect with an estimate of β_3 equal to 1.29 (standard error 0.29). In Specification 2 we decompose the effect of diversity into the effect of the increased share of foreign-born and the effect of increased diversity among the foreign-born. Both measures have a positive and significant effect on wages, but the effect from increased share of foreign-born is much more precisely estimated at 0.58 with standard error 0.1. Increased diversity of foreigners has an impact significant at the 10-per-cent level only. Specifications 3 and 4 estimate the effect of diversity on the average income of white US-born males with 40 to 50 years of age, which includes returns to capital and entrepreneurship. As long as diversity acts as a local production externality, its effect should also affect these returns. Reassuringly diversity has a positive and significant effect on personal income as well, and even larger than on wage income. According to the estimates in Specification 3, increasing the diversity index by 10 percent would cause an increase in average personal income of US-born by 15 percent. Decomposing the effect of diversity into the effects of the share of foreign-born and of diversity among foreign-born, the former turns out to be the key component while the latter effect is positive but not significant. An increase in foreign-born by 10 percent would increase average personal income of US-born by 8.2 percent.

Table 3
Basic Panel Wage Regression

Explanatory Variables:	**1: Dependent Variable: ln(Wage)**	**2: Dependent Variable: ln(Wage)**	**3: Dependent Variable: ln(Income)**	**4: Dependent Variable: ln(Income)**
Average Schooling	0.10** (0.01)	0.10** (0.01)	0.07* (0.01)	0.07* (0.01)
Ln(Employment)	0.02 (0.02)	0.01 (0.02)	0.14* (0.03)	0.10** (0.03)
Diversity Index	1.29** (0.29)		1.55** (0.70)	
Share of Foreign Born		0.58** (0.11)		0.82* (0.27)
Diversity Index Among Foreign Born		0.14* (0.08)		0.05 (0.10)
City Fixed Effects	Yes	Yes	Yes	Yes
Time Fixed Effects	Yes	Yes	Yes	Yes
R^2	0.99	0.99	0.99	0.99
Observations	320	320	320	320

I and II: Dependent Variable: ln average yearly wage of white, U.S. Born, males 40-50 years in 1990 U.S. $.
III and IV: Dependent Variable: ln average yearly income of white, U.S. Born, males in 1990 U.S. $.
** significant at 5%, * significant at 10%
In Parenthesis: Heteroskedasticity-Robust Standard Errors.

Vis a vis the positive and significant effect of diversity on wages, it is then crucial for identification to measure the impact of diversity on rents. Table 4 reports the results from the rent regression (15). Specification 1 and 2 control only for population density plus city and year fixed effects. Specification 3 and 4 control for personal income too. Again we estimate the effect of overall diversity (Specifications 1 and 3) and then we decompose it into the effects of the share of foreign-born and of diversity of foreign-born (Specifications 2 and 4). Considering the impact of the share of foreign-born, which turns out to be the most important component of diversity, when we do not control for personal income (Specification 2), an increase of the share of foreign-born by 10 percent is associated with an increase in rents for US-born close to 11 percent. As reported in Table 3, though, an increase in diversity is associated with an increase in personal income so that the effect on rents may be a consequence of higher average income in the city without any independent additional effect. However, when we control for average income of the US-born group (Specification 4), we still have a positive and significant effect on US-born rents although half the size of the one estimated in Specification 1. An increase of the share of foreign-born by 10 percent would increase price of housing by 5.3 percent, even after controlling for the fact that higher diversity is associated with higher income, and 1 percent higher income generates 0.6 percent higher rents.

Table 4
Basic Panel Rent Regression

Explanatory Variables:	**1: Dependent Variable: ln(Rent)**	**2: Dependent Variable: ln(Rent)**	**3: Dependent Variable: ln(Rent)**	**4: Dependent Variable: ln(Rent)**
Ln(Income per Capita)			0.67** (0.08)	0.66** (0.08)
Ln(Population)	0.10** (0.04)	0.02 (0.04)	0.03 (0.04)	0.06 (0.04)
Diversity Index	1.80** (0.60)		0.95** (0.50)	
Share of Foreign Born		1.06** (0.27)		0.53** (0.20)
Diversity Index Among Foreign Born		0.11 (0.16)		0.16 (0.13)
City Fixed Effects	Yes	Yes	Yes	Yes
Time Fixed Effects	Yes	Yes	Yes	Yes
R^2	0.97	0.99	0.98	0.99
Observations	320	320	320	320

Dependent Variable: ln average monthly rent per room paid by white, U.S. Born, expressed in 1990 U.S. $.
** significant at 5%, * significant at 10%
In Parenthesis Heteroskedasticity Robust Standard Errors.

To sum up, diversity has *positive and highly significant correlations* with both wage ($\beta_3 > 0$) and land rent ($\gamma_3 > 0$). According to (13), such positive correlations can be interpreted as consistent with a dominant production amenity effect of diversity. To gain further insight on this result, the rest of the paper is devoted to two tasks. First, in Section 5.2 we check whether those positive correlations survive the inclusion of several additional controls. Second, in Section 5.3 we tackle the issue of endogeneity raised at the end of Section 4.2. In particular, we try to assess the causal direction of those correlations by instrumental variables techniques.

Before doing that, however, let us check another correlation that may reinforce our interpretation that the positive effects on wages and rents are the equilibrium result of a dominant production amenity. The theoretical model makes clear (see (6)) that, in the presence of a production amenity, labor demand would shift up in cities where diversity increased. Table 5 reports the correlation between changes in diversity and changes in employment as well as population of US cities between 1970 and 1990. If the labor supply curve had shifted up with labor demand unchanged, that would have caused the observed increase in wages but this would have been associated with a decrease in employment. On the contrary, Table 5 shows mildly positive effects of diversity on employment and population, not significant the former and significant the latter. Such results, therefore, point to some dominant upward shift of labor demand as expected in the presence of a dominant production amenity.

Table 5
Correlation between Diversity and Employment/Population

	Index of Diversity	**City Fixed Effects**	**Time fixed effects**	**R^2**	**Observations**
Ln(Employment)	0.72 (1.12)	Yes	Yes	0.97	320
Ln(Population)	1.70* (1.02)	Yes	Yes	0.97	320

** significant at 5%, * significant at 10%
In Parenthesis Heteroskedasticity Robust Standard Errors.

5.2 Check of Robustness

Our basic specifications for the wage and rent regressions omit several variables that, in principle, could affect both the degree of diversity and local externalities. In so far as they change over time, the impacts of such omitted variables are not captured by the city fixed effects.

This section is devoted to testing whether the estimated effects of diversity are robust to the inclusion of omitted variables. While the list of potential controls is never complete, we include here some important ones for which one can think of plausible stories that would lead to the estimated correlations. Table 6 reports the estimates of the coefficients of the diversity index, the share of foreign-born, and foreign-born diversity in the wage equation as we include additional controls, one at a time and together. Table 7 presents analogous results for the rent regression.

In addition, our theoretical model shows that equilibrium wages and rents are simultaneously determined. This suggests that there may be correlation between the unobservable idiosyncratic shocks to wages, ε_{ct}, and rents, e_{ct}. To deal with this potential source of inefficiency in OLS estimations, Table 8 reports the coefficients of the diversity index or the share of foreign-born in the wage and rent equations when simultaneously estimated by SUR.

5.2.1 Skills' Complementarity/Externality from Foreign-Born

The positive effect of the foreign-born on the US-born wage could simply be a result of the foreigners' measurable skills. If the foreign-born had higher (or lower) schooling achievements than US-born, then, through some complementarity or externality effects, that could increase the wages of the US-born independently from any role of diversity. Ciccone and Peri (2002) find a significant complementarity between human capital and labor in US cities and Moretti (2003) finds significant externalities from schooling. Thus, there is some ground to suspect that we might be attributing to diversity an effect more simply due to the observable levels of schooling of the foreign-born.

Specifications (2) in Tables 6 and 7 include the average years of schooling of the foreign-born as a control variable in the wage and rent regressions respectively. The effect of diversity is still significant and positive in both cases. Interestingly, the effect (not reported) of average schooling of the foreign-born on the wages of the US-born is not significant, while it is small and positive on US-born rents. When decomposing the overall diversity effect (column 2 and 3 in the tables) we find a significant and positive effect of the share of foreign-born on both rents and wages, while the diversity of foreigners has significant positive impact on wages but not on rents. As in most specifications, in this case too the estimated effect implies that an increase of 10 percent in the share of foreign-born increases average US-born wages as well as rents by 5.8.

5.2.2 Unobserved Shocks to Productivity and Amenities

Another plausible reason to find positive correlations of diversity with wages and rents may be that migration into the city responds to positive shocks to productivity and local amenities. In so far as we do not observe these shocks, we are omitting the common underlying cause of increased wages, increased rents, and increased diversity. To address this issue we use two strategies. The first strategy, which we postpone to Section 5.3, tries to identify a variable correlated (or more correlated) with the share of foreign-born but not otherwise correlated with productivity or amenities. Then, it uses such variable as instrument for the estimation.

The second strategy, pursued here, exploits the fact that, if shocks to productivity attract workers into a city, this should work for US-born as well as for foreign-born workers. Therefore, if we included the share of US-born citizens coming from out of state (i.e., born in a different state than the state of the city of residence) in the wage and rent regressions, such variable should be correlated with productivity and amenities shocks too, and therefore its inclusion should decrease significantly the estimated coefficients β_3 and γ_3. Moreover, we should find a significant positive correlation between wages as well as rents and the share of people born out of state. Specification (3) in Tables 6 and 7 include the share of US-born citizens who were born out of state. The coefficients on this variable (not reported) are not significant in either regression, while the effects of diversity and of the share of foreign-born on wages and rents are still significantly positive and virtually unchanged. These results suggest that the presence of the foreign-born does not simply signal that cities have some unobserved positive characteristics that would lure both foreign- and US-born workers.

5.2.3 Tolerance for Diversity

Some sociologists have advanced the hypothesis that environments that are tolerant towards diversity are more productive and more pleasant to live in: they are more open to contribution of ideas from different groups and they are more amenable. Along these lines Richard Florida (see, e.g., Florida, 2002) has argued that cities where the number of artists and bohemian professionals is larger are more innovative in high tech sectors. It is quite likely that part of our correlations may actually depend on this good attitude of cities towards diversity. However, to show that there is nonetheless something specific to the presence of foreign-born in generating the positive correlation of diversity with wages and rents, we include another measure of diversity in our regressions.

The share of US-born people identifying themselves as 'non-white', is entered in specification (4) of Tables 6 and 7. Since we consider only US-born people, such index essentially captures the white-black composition of a city. The coefficients (not reported) on this variable turn out to be positive in the wage regression (0.20) and negative in the rent regression (-0.22). According to (13), this would suggest a dominant consumption disamenity effect of the share of non-white US-born. While it might be tempting to read this result as evidence of the aversion of white US-born against living close to large non-white communities, the standard errors (in both cases around 0.2) make the estimated coefficients not significant.

As to the coefficients of the diversity index and of the share of foreign-born, they are still positive, significant (except in one case for the rent regression), and similar to previous estimates. Thus, independently from the impact of diversity along the ethnic dimension, diversity in terms of the country of birth maintains its own importance and specificity.

5.2.4 Quality of Local Public Goods

Several public services in US cities are supplied by local governments. Public schools, public health care, and public security are all desirable local services. Therefore, cities where their quality has improved in the period of observation may have experienced both an increase in the share of foreign-born (possibly larger users of these services) and a rise in property values. From the County and City Databook we have gathered data on the spending of local government per person in a city and on its breakdown across different categories such as education, health care, and security (local police). Specification (5) includes overall spending by local government whereas Specification (6) includes spending on education, a very important determinant of the quality of schools, which in turn largely affects property value.

Once more, while the effect of public spending per person on rents (not reported) is positive in both specifications, the inclusion of the controls does not change the effect of diversity. In particular, the coefficient on the share of foreign-born is around 0.6 (standard error 0.17) in the wage regression and 0.53 (standard error 0.24) in the rent regression. As a final and most conservative check, Specification (7) includes together all the controls included separately in specifications from 2 to 6. Reassuringly, the coefficient of the share of foreign-born is still positive, very stable, and significant in both regressions. The coefficient of the diversity index is also positive, very stable, and significant in the wage regression while it turns out not significant in the rent regression.[9]

5.2.5 Outliers and State by Year Fixed Effects

In the last two specifications of Tables 6 and 7 we try to push our data as far as they can go. Specification (8) estimates the wage and rent regressions excluding the three states with the highest shares of foreign-born, namely California, New York and Florida. The aim is to check whether few highly diverse cities in those states generate the correlations of diversity with wages and rents estimated for the whole data set. This is not the case. In the wage regression the coefficient of diversity decreases somewhat but remains both positive and significant. In the rent equation the coefficient of diversity becomes much larger but also much less precisely estimated. In general, however, there is no evidence that in the long run the effect of diversity is different for high immigration states than for low immigration states.

In Specification (9), rather than the panel with city and year dummies, we use instead the differences between 1990 and 1970 of the basic variables. We also include state fixed effects to control for differences in state-specific growth rates of wages and rents. In so doing we identify the effects of diversity on wages and rents through the variation across cities within states. Indeed, the inclusion of state fixed effects in the difference regression is equivalent to the inclusion of state by time fixed effects in the panel regression. This is an extremely conservative specification as we are probably eliminating a lot of the variation needed to identify the results and we are estimating 48 dummies using 160 observations. Remarkably, the positive effect of diversity on productivity still stands and its point estimate is similar to those of previous specifications. The effect of diversity on rents, however, while still positive, is no longer significant.

[9] Some authors (see, e.g., Sivitanidou and Wheaton, 1992) have argued that also the institutional constraints on land use ('zoning') can affect land values. Thus, higher property values may be associated with more efficient institutional constraints in the presence of market failures. Also this effect should be captured by our local public goods measures.

Table 6
Wage Regression: Robustness Checks

Specification	**1 Coefficient on the Diversity Index**	**2 Coefficient on the Share of Foreign Born**	**3 Coefficient on Diversity Index Among Foreign Born**
Specification:			
(1) Basic	1.29** (0.29)	0.58** (0.11)	0.14* (0.08)
(2) Including schooling of Foreign Born	1.25** (0.38)	0.58** (0.16)	0.14* (0.09)
(3) Including share of Out of State Born	1.34** (0.38)	0.61** (0.16)	0.09 (0.09)
(4) Including share of non Whites	1.38** (0.40)	0.68** (0.18)	0.13 (0.10)
(5) Including Public Spending on local Services per capita	1.28** (0.38)	0.62** (0.17)	0.16** (0.08)
(6) Including Public Spending in Education per capita	1.24** (0.38)	0.61** (0.16)	0.15* (0.08)
(7) Including All of the Above	1.40** (0.40)	0.74** (0.18)	0.09 (0.08)
(8) Basic without CA, FL, NY	0.97** (0.50)	0.30 (0.27)	0.21** (0.10)
(9) in Changes 1990-1970 with State-Fixed Effects	1.05** (0.33)	0.64** (0.18)	0.03 (0.10)

Dependent Variable: ln average yearly wage to white, U.S. Born, 40-50 years old expressed in 1990 U.S. $.

(1) Basic: Specification from Table 3 Column I (for coefficient 1) and Column II (for coefficients 2 and 3)
(2) Includes average years of schooling of foreign born
(3) Includes the share of U.S. born outside the state in which they live
(4) Includes the share of non-white people in working age
(5) Include the Spending per capita on local government services.
(6) Includes the Spending in Education per Capita
(7) Includes all the variables in (1)-(6) together as controls
(8) Excluding from the regression MSAs in the biggest immigrations states (CA, FL, NY)
(9) Regression in Changes including 49 State Fixed-Effects

** significant at 5%, * significant at 10%
In Parenthesis Heteroskedasticity Robust Standard Errors.

Table 7
Rent Regression: Robustness Checks

Specification	**1 Coefficient on the Diversity Index**	**2 Coefficient on the Share of Foreign Born**	**3 Coefficient on Diversity Index Among Foreign Born**
Specification:			
(1) Basic	0.95** (0.50)	0.53** (0.20)	0.11 (0.16)
(2) Including schooling of Foreign Born	0.92* (0.49)	0.58** (0.22)	0.17 (0.14)
(3) Including share of Out of State Born	0.86* (0.50)	0.52** (0.24)	0.16 (0.15)
(4) Including share of non Whites	0.71 (0.50)	0.44** (0.25)	0.14 (0.13)
(5) Including Public Spending on local Services per capita	0.89* (0.50)	0.53** (0.24)	0.18 (0.14)
(6) Including Public Spending in Education per capita	0.94** (0.50)	0.53** (0.24)	0.14 (0.14)
(7) Including All of the Above	0.72 (0.50)	0.53** (0.24)	0.15 (0.14)
(8) Basic without CA, FL, NY	2.99** (1.19)	0.74* (0.39)	0.10 (0.10)
(9) in Changes 1990-1970 with State-Fixed Effects	0.13 (0.42)	0.03 (0.23)	0.15 (0.17)

Dependent Variable: ln average monthly Rent paid by white, U.S. Born, expressed in 1990 U.S. $.

(1) Basic: Specification from Table 4 Column II (for coefficient 1) and Column IV (for coefficients 2 and 3)
(2) Includes average years of schooling of foreign born
(3) Includes the share of U.S. born outside the state in which they live
(4) Includes the share of non-white people in working age
(5) Include the Spending per capita on local government services.
(6) Includes the Spending in Education per Capita
(7) Includes all the variables in (1)-(6) together as controls
(8) Excluding from the regression MSAs in the biggest immigrations states (CA, FL, NY)
(9) Regression in Changes including 49 State Fixed-Effects

** significant at 5%, * significant at 10%
In Parenthesis Heteroskedasticity Robust Standard Errors.

5.2.6 Efficient Estimation: SUR

The theoretical model shows that in equilibrium wages and rents are simultaneously determined (see equations (11) and (12)). This implies that there may be correlation between the unobservable idiosyncratic shocks to wages, ε_{ct}, and rents, e_{ct}. If so, we could increase the efficiency of our estimates by explicitly accounting for such correlation and estimating the wage and rent equations as a system through SUR. While OLS estimates are still consistent and unbiased even when ε_{ct} and e_{ct} are correlated, SUR estimates are more efficient.

Table 8 reports the coefficients on the diversity index or the share of foreign-born in the wage and rent equations when estimated simultaneously by SUR. Specifications (1) through (9) are identical to (1) through (9) in Tables 6 and 7. In general, the estimates of the coefficients on the diversity index are somewhat larger when using SUR in both the wage and the rent equations. When we use the share of foreign-born instead, the estimated coefficients are similar to the OLS results for wages but somewhat higher for rents. Indeed, while the elasticity of wages to the share of foreign-born is between 0.6 and 0.7 in most specifications (just as in Table 6), the elasticity of rents is between 0.6 and 1.1 (versus $0.6 - 0.8$ in Table 7). The only case in which the effect of diversity on wages is only borderline significant is when we exclude the large immigration states (just as in the OLS estimates). As to rents, the effect of diversity is not significant only when we introduce state-specific trends (again, as in the OLS estimates). Therefore, SUR estimates strongly confirm the qualitative and quantitative results obtained under OLS.

Table 8
SUR Estimation of the system of Wage and Rent equations

Dependent Variable	**Wage**		**Rents**	
Coefficient Estimate on:	**Diversity Index**	**Share of Foreign Born**	**Diversity Index**	**Share of Foreign Born**
Specification:				
(1) Basic	1.78** (0.42)	0.56** (0.11)	1.24** (0.27)	1.05** (0.19)
(2) Including schooling of Foreign Born	1.77** (0.40)	0.56** (0.11)	1.24** (0.26)	1.09** (0.18)
(3) Including share of Out of State Born	1.80** (0.42)	0.61** (0.11)	1.30** (0.26)	1.07** (0.19)
(4) Including share of non Whites	1.41** (0.43)	0.67** (0.13)	1.36** (0.28)	0.88* (0.20)
(5) Including Public Spending on local Services per capita	1.77** (0.42)	0.59** (0.11)	1.27** (0.26)	1.05** (0.18)
(6) Including Public Spending in Education per capita	1.80** (0.42)	0.58** (0.11)	1.23** (0.26)	1.06** (0.19)
(7) Including All of the Above	1.13** (0.27)	0.58** (0.13)	0.79** (0.35)	0.56** (0.17)
(8)Basic without CA, FL, NY	1.09* (0.60)	0.34 (0.21)	3.50** (0.90)	1.21* (0.30)
(9) in Changes 1990-1970 with State-Fixed Effects	1.05** (0.34)	0.64** (0.16)	0.11 (0.19)	0.03 (0.19)

System estimation, Dependent variables :
ln average yearly wage to white, U.S. Born, 40-50 years old expressed in 1990 U.S. $.
ln average monthly Rent paid by white, U.S. Born, expressed in 1990 U.S. $.

(1)Basic: Specification
(2) Includes average years of schooling of foreign born
(3) Includes the share of U.S. born outside the state in which they live
(4) Includes the share of non-white people in working age
(5) Include the Spending per capita on local government services.
(6) Includes the Spending in Education per Capita
(7) Includes all the variables in (1)-(6) together as controls
(8) Excluding from the regression MSAs in the biggest immigrations states (CA, FL, NY)
(9) Regression in Changes including 49 State Fixed-Effects

** significant at 5%, * significant at 10%

5.2.7 Overview of Robustness Checks

To summarize the results of our robustness checks, in Tables 6, 7 and 8 most wage and rent regressions yield positive and significant coefficients for both the diversity index and the share of foreign-born. In only one case (Specification (8)) the OLS estimates of the single wage regression produce an insignificant effect of the share of foreign-born. In only three cases the coefficient of diversity (Specifications (4), (7), (9)) or the coefficient of the share of foreign-born (Specification (9)) are not significant in the OLS estimates of the single rent regression. As to SUR estimation, in Table 8 only two specifications find an insignificant effect of diversity either on wages (Specification (8)) or on rents (Specification (9)). We do not find any specification such that the coefficients of the same variable are simultaneously not significant in both wage and rent regressions.

Therefore, our identification (13) allows us to conclude that *all specifications support the hypothesis of a production amenity value of diversity.* Indeed, by using Figure 3, in subsection 4.2 we have already argued that, when the coefficients of our diversity measures are positive and significant in both the wage and the rent regressions, this is evidence of a dominant production amenity (see point A'). In addition, it is readily established that a significant positive impact of diversity on wages and no significant impact on rents is consistent only with a production amenity and a consumption disamenity (see point B). Analogously, a significant positive impact of diversity on rents and no significant impact on wages is consistent only with both a production amenity and a consumption amenity (see point C).

5.3 Endogeneity and Instrumental Variables

Short of a randomized experiment in which diversity across cities is changed exogenously and randomly, we cannot rest assured that our correlations reveal any causal link from diversity to wages and rents. Nonetheless, some steps towards tackling such endogeneity problem can be taken by instrumental variables (IV) estimation. Our instruments should be correlated with the change in diversity of cities in the 1970-1990 and not otherwise correlated with changes in wages and rents. We propose two instruments satisfying the foregoing properties. Both exploit the fact that, presumably exogenously from the characteristics of any single city, the overall immigration to the US increased significantly between 1970 and 1990.

5.3.1 Ports of entry

To construct the first instrumental variable, we build on the fact that immigrants tend to enter the US through few 'ports of entry'. Each year the US Office of Tourism publishes the percentage of inbound travellers by port of entry. Looking at the data for the eighties, we see that the three main gateways were New York, Miami, and Los Angeles. Through the airports and ports of these cities about 30 percent of the foreign (immigrant and non-immigrant) travellers entered the US. Accordingly, foreign-born immigrants from countries other than Canada or Mexico were quite likely to go through one of those cities. Moreover, due to networks, costs of travelling, and costs of spreading information, such immigrants were more likely to settle down in cities closer to those gateways. Therefore, cities at a smaller distance from those ports of entry were more likely to receive foreign-born immigrants during the 1970-1990 period. A similar argument can be made for Canadian and Mexican immigrants. For them it seems reasonable to assume that the US borders with their own countries constitute their natural port of entry. Thus, as before, cities at a smaller distance from the borders were more likely to receive Canadian and Mexican immigrants during the 1970-1990 period.

Such considerations suggest the use of the overall distance of a city from the main 'ports of entry' (New York, Miami, Los Angeles, and the US borders with Canada and Mexico) to instrument its share of foreign-born or its diversity index. On the one hand, such distance should be negatively correlated with diversity. As stated above, cities closer to the ports of entry experienced larger inflows of foreign-born in the 1970-1990 period. On the other hand, during the same period the distance of a city from the ports of entry should also be little correlated with changes in its wages and rents.

This strategy is open to an obvious critique. If, for example, the three cities (New York, Miami, and Los Angeles) had experienced above average growth in the considered time period, this could have had positive spillover effects on nearby cities. As a result, the distance of a city from them would be negatively correlated with the increases in wages and rents due to the spillovers. To see whether this indeed happened in our data set, we have calculated employment growth for the three cities in the period of observation. It turns out that in each of those three cities employment growth was actually lower than average (respectively, +45 percent in Los Angeles, +42 percent in Miami, and a mere +7 percent in New York, against an average +48 percent). Also population growth was below average in both New York and Los Angeles, while it was above average in Miami. Therefore, overall the three cities did not really grow faster than other cities. However, since Miami still exhibited better performance than the other two ports of entry, we have also used only distances from New York and Los Angeles as instruments: the results are virtually unchanged.

Table 9
IV Estimation, Instrument: Distance from port of Entry.
Wage Regression

Dependent Variable : Δln(Wage) 1970-1990	**1 Diversity Index**	**2 Share of Foreign Born**	**3 Diversity Index**	**4 Share of Foreign Born**	**5 Non-Coastal Cities**	**6 Non-Coastal Cities**
ΔSchooling	0.11** (0.01)	0.10** (0.01)	0.10** (0.02)	0.11** (0.02)	0.11** (0.02)	0.10** (0.02)
Δln(Empl)	0.02 (0.02)	0.01 (0.02)	0.07 (0.04)	0.05 (0.04)	0.04 (0.04)	0.07 (0.05)
Δ(Foreign Born)		0.72* (0.12)		0.69** (0.20)		1.60** (0.47)
Δ(Diversity)	1.55** (0.47)		1.23** (0.42)		4.70* (2.40)	
State Fixed Effects	No	No	Yes	Yes	No	No
R^2	0.35	0.35	0.65	0.66	0.30	0.30
Observations	160	160	160	160	144	144
First Stage Estimation						
Ln(Distance from LA)	-0.038** (0.004)	-0.015** (0.001)	-0.038** (0.004)	-0.015** (0.001)	-0.034** (0.006)	-0.007** (0.001)
Ln(Distance from NY)	-0.004** (0.002)	-0.004** (0.001)	-0.004** (0.002)	-0.004** (0.001)	-0.003 (0.003)	-0.004** (0.001)
Ln(Distance from Miami)	-0.023** (0.003)	-0.012* (0.001)	-0.023** (0.003)	-0.012* (0.001)	-0.023** (0.007)	-0.0005 (0.0005)
Ln(Distance from Border)	-0.002** (0.001)	-0.001 (0.001)	-0.002** (0.001)	-0.001 (0.001)	-0.004** (0.002)	-0.001** (0.0005)
R^2	0.55	0.54	0.55	0.54	0.36	0.30

Dependent Variable: Change between 1970 and 1990 in ln average yearly wage of white, U.S. Born, 40-50 years, expressed in 1990 U.S. $.
** significant at 5%, * significant at 10%
In Parenthesis Heteroskedasticity Robust Standard Errors.

Table 10
IV Estimation, Instrument: Distance from Port of Entry.
Rent Regression

Dependent Variable : **Δln(Rent) 1970-1990**	**1** **Diversity Index**	**2** **Share of Foreign Born**	**3** **Diversity Index**	**4** **Share of Foreign Born**	**5** **Non-Coastal Cities**	**6** **Non-Coastal Cities**
Δln(Income)	0.65** (0.10)	0.65** (0.10)	0.47** (0.12)	0.46** (0.12)	0.39* (0.17)	0.29** (0.11)
Δln(Pop)	0.03 (0.04)	0.03 (0.04)	0.04 (0.08)	0.03 (0.08)	0.02 (0.14)	0.02 (0.10)
Δ(Foreign Born)		1.05** (0.50)		0.20 (0.33)		1.40** (0.68)
Δ(Diversity)	1.60 (1.00)		0.17 (0.48)		5.90** (1.80)	
State Fixed Effects	No	No	Yes	Yes	No	No
R^2	0.35	0.36	0.73	0.73	0.14	0.18
Observations	160	160	160	160	144	144
First Stage Estimation						
Ln(Distance from LA)	-0.038** (0.004)	-0.015** (0.001)	-0.038** (0.004)	-0.015** (0.001)	-0.034** (0.006)	-0.007** (0.001)
Ln(Distance from NY)	-0.004** (0.002)	-0.004** (0.001)	-0.004** (0.002)	-0.004** (0.001)	-0.003 (0.003)	-0.004** (0.001)
Ln(Distance from Miami)	-0.023** (0.003)	-0.012* (0.001)	-0.023** (0.003)	-0.012* (0.001)	-0.023** (0.007)	-0.0005 (0.0005)
Ln(Distance from Border)	-0.002** (0.001)	-0.001 (0.001)	-0.002** (0.001)	-0.001 (0.001)	-0.004** (0.002)	-0.001** (0.0005)
R^2	0.55	0.54	0.55	0.54	0.36	0.30

Dependent Variable: Change between 1970 and 1990 in ln average monthly rent paid by white, U.S. Born, expressed in 1990 U.S. $.
** significant at 5%, * significant at 10%
In Parenthesis Heteroskedasticity Robust Standard Errors.

Table 9 and 10 report the first and second stage estimates of the described IV regression. Columns 1 and 2 show the basic specification; columns 3 and 4 include 48 state fixed-effects; columns 5 and 6 exclude all the coastal cities from the regression to make sure that our results are not simply driven by the difference in shares of foreigners and productivity between the coast and the inland. The first stage regressions confirm that our instruments are excellent and explain about 50 percent of the variation of diversity across cities. Farther from the ports of entry diversity is significantly lower.

Considering specifications 1 and 2, we find that the OLS results are confirmed, the effect of the share of foreign-born on wages and rents across cities is positive and significant. Moreover, the IV estimates are somewhat higher than the OLS ones, so that we are reassured that endogeneity did not cause a significant OLS bias. Again, the coefficients on the wage regression confirm that an increase in foreign-born by 1 percent is associated with 0.7 percent increase in wages. As for the rents, the impact is around 1 percent. For the wage regressions we obtain a positive significant effect of diversity also when controlling for 48 state fixed effects (specifications 2 and 3) and when we eliminate coastal cities (specifications 5 and 6). These last two specifications have quite large standard errors, however, but certainly reinforce our thesis that foreign-born have a positive effect also in non-coastal cities. As to the rent regressions, the share of foreigners has a positive and significant effect in specification 2 and 6. When we include state dummies, the effect of foreign-born is not significant any longer (but still positive) as in Table 7. Again, the estimates on non-coastal cities have large standard errors but the coefficient estimate on the share of foreign-born is robustly positive and close to one. Differently, the coefficient of the diversity index is positive and significant only for non-coastal cities.

5.3.2 Shift-Share Methodology

A second instrumental variable, independent of idiosyncratic city shocks to wages and rents, could be constructed by adopting the 'shift-share methodology' used by Card (2001) and, more recently, applied also by Saiz (2003b) to migration in MSAs. Immigrants tend to settle where other immigrants from the same country already reside. Therefore, we can use the share of residents of an MSA in 1970 born in each country to attribute to each group the growth rate of its share within the whole US population in the 1970-1990. In so doing we compute the predicted composition of the city based on its 1970 composition and attributing to each group the average growth rate of its share in the US population. Once we have constructed these 'predicted' shares for 1990 we can calculate the 'attributed' diversity index and the 'attributed' share of foreign-born for each city in 1990.

Let us use the notation introduced in section 3.1 where $(CoB_j^c)_t$ labels the share of people born in country j among the residents of city c in year t. Hence, $(CoB_j)_t = \sum_c (CoB_j^c)_t$ is the share of people born in country j among US residents in year t so that between 1970 and 1990 its growth rate is:

$$(g_j)_{1970-90} = [(CoB_j)_{1990} - (CoB_j)_{1970}]/(CoB_j)_{1970} \qquad (16)$$

This allows us to calculate the 'attributed' share of people born in country j and residing in city c in 1990 as:

$$(\widehat{CoB_j^c})_{1990} = (CoB_j^c)_{1970} \bullet [1 + (g_j)_{1970-90}] \qquad (17)$$

The attributed share of foreign-born and the attributed diversity index can evaluated accordingly. In particular, the latter equals:

$$div(\widehat{CoB})_{c1990} = 1 - \sum_j (\widehat{CoB_j^c})^2_{1990} \qquad (18)$$

As the attributed values for each city in 1990 are built using the city's shares in 1970 and the national growth rates of groups from 1970 to 1990, they are independent from any city-specific event that period. Thus, being orthogonal to city-specific shocks, they can be used as instruments for the actual values.

Table 11 and 12 present the results of the IV estimation of the wage and rent regressions using the shift-share instruments. Unfortunately in terms of perfect comparability with previous results, some adjustments in the grouping of countries of birth is unavoidable. The reason is that, as we input the shares in 1990 based on the initial shares in 1970 and the national growth of groups, we need to identify the same groups across census years. This is achieved by allocating more than one country of birth to the same group. In so doing, we follow the classification adopted by Card (2001), Table 5, as described in the data appendix.

Table 11
IV Estimation, Instrument: Shift-Share constructed Diversity.
Wage Regression

Dependent Variable : Δln(Wage) 1970-1990	**1 OLS Diversity Index**	**2 OLS, Share of Foreign Born**	**3 IV, Diversity Index**	**4 IV, Share of Foreign Born**	**5 IV Without CA-FL-NY**	**6 IV Without CA-FL-NY**
ΔSchooling	0.11** (0.01)	0.10** (0.01)	0.11** (0.01)	0.11** (0.01)	0.10** (0.02)	0.10** (0.01)
Δln(Empl)	0.02 (0.02)	0.01 (0.02)	0.04 (0.03)	0.04 (0.03)	0.03 (0.03)	0.04 (0.03)
Δ(Foreign Born)		0.51** (0.10)		0.30 (0.41)		0.22 (0.22)
Δ(Diversity)	1.27** (0.27)		0.95** (0.50)		0.92 (0.65)	
R^2	0.36	0.35	0.35	0.24	0.34	0.22
Observations	160	160	160	160	145	145
First Stage Regression, for the IV estimation						
Shift-Share Constructed Diversity	n.a.	n.a.	0.51** (0.05)	0.32** (0.03)	0.21** (0.04)	0.23** (0.03)
R^2	n.a.	n.a.	0.34	0.37	0.15	0.31

Dependent Variable: Change between 1970 and 1990 in ln average yearly wage of white, U.S. Born, 40-50 years, expressed in 1990 U.S. $.
Instrumental Variable: Imputed change in diversity index and share of foreign born, using the shift-share method.
** significant at 5%, * significant at 10%
In Parenthesis Heteroskedasticity Robust Standard Errors.

Table 12
IV Estimation, Instrument: Shift-Share constructed Diversity.
Rent Regression

Dependent Variable : Δln(Rent) 1970-1990	**1 OLS, Diversity Index**	**2 OLS, Share of Foreign Born**	**3 IV, Diversity Index**	**4 Share of Foreign Born**	**5 Without CA-FL-NY**	**6 Without CA-FL-NY**
Δln(Population)	0.03 (0.04)	0.06 (0.04)	0.04 (0.04)	0.08** (0.04)	0.04 (0.06)	0.09 (0.06)
Δln(Income)	0.67* (0.09)	0.64* (0.09)	0.61* (0.10)	0.59** (0.09)	0.48** (0.09)	0.51* (0.08)
Δ(Foreign Born)		0.58** (0.29)		0.98** (0.36)		0.74 (0.50)
Δ(Diversity)	1.10* (0.70)		2.60** (1.02)		4.21** (1.60)	
R^2	0.38	0.37	0.33	0.36	0.28	0.28
Observations	160	160	160	160	145	145
First Stage						
Shift-Share Constructed Diversity	n.a.	n.a.	0.51** (0.05)	0.32** (0.03)	0.21** (0.04)	0.23** (0.03)
R^2	n.a.	n.a.	0.34	0.37	0.15	0.31

Dependent Variable: Change between 1970 and 1990 in ln average yearly wage of white, U.S. Born, 40-50 years, expressed in 1990 U.S. $.
Instrumental Variable: Imputed change in diversity index and share of foreign born, using the shift-share method.
** significant at 5%, * significant at 10%
In Parenthesis Heteroskedasticity Robust Standard Errors.

In Tables 11 and 12, columns 1 and 2 report the OLS estimates of the basic specifications for the two measures of diversity. The point estimates of the OLS specification are very similar to the previous estimates (respectively Table 3 Columns 1 and 2 and Table 4 columns 3 and 4) confirming that the reclassification of groups has only negligible effects. The first stage regressions shows that the attributed and the actual changes are positively correlated, with the former explaining 30-50 percent of the variation of the latter when all states are included.

The IV-estimated effect of diversity on wages is reported in column 3 and 4 of Table 11. The effect is still positive and significant when diversity is measured using the diversity index (specification 3). When we include only the share of foreign-born as measure of diversity, however, its coefficient is still positive but the standard error increases and the estimate is not significant. Similarly when we exclude the high-immigration states, the effect of diversity is estimated to be positive but not significant. However, the main problem when we exclude California, Florida, and New York is that the instruments lose much of their explanatory power (only 20 percent of the variance of the endogenous variable is explained by the instrument). Therefore, insignificance is mostly driven by large standard errors rather than by evidence of endogeneity bias (i.e., changes in point estimates).

In Table 12 the rent regression exhibits a similar qualitative pattern but sharper results. Using the shift-share instruments both the diversity index and the share of foreign-born have a positive and significant effect (specifications 3 and 4). Also in this regression, when we exclude California, Florida, and New York, the standard errors increase significantly. However, the point estimates of the effect of diversity are still firmly in the positive range. Somewhat surprising and possibly driven by some outliers is the very large (and imprecisely estimated) effect of diversity on rents in specification 5. As similar results emerged in Table 10 when coastal cities were excluded.

5.3.3 Overview of Instrumental Variables

As the theoretical model makes clear, endogeneity is a potential problem for our results. We have tackled such problem by IV estimation using two different sets of instruments.

When distance from the ports of entry is used, the estimated coefficients of both diversity measures are always positive and significant in the wage regressions. Differently, in the rent regressions they are significant, and positive, in one case only, namely, when we use the diversity index while excluding non-coastal cities.

When the shift-share approach is adopted instead, we have symmetric results. In the rent regressions diversity measures have always positive and significant coefficients except in one case. The exception is when we use the share of foreign-born while excluding California, Florida, and New York, in which case the estimated coefficient is not significant. Differently, in the wage regressions the estimated coefficients of the diversity measures are significant, and positive, in one case only, namely, when we use the diversity index without excluding California, Florida, and New York.

All in all, these results mirror the results of the OLS regressions. In particular, in the IV regressions we find only one specification such that the coefficients of the same variable are simultaneously not significant in both wage and rent equations. This is never the case in the OLS regressions. Thus, on the basis of the discussion in subsection 5.2.7, we can conclude that our data support the hypothesis of a production amenity value of diversity with *causation running from diversity to the location decisions of firms and workers.*

6 Discussion and Conclusions

We have developed a theoretical model of an urban system of open cities in which cultural diversity affects production and consumption as an externality. In principle, the external effects of diversity can be positive ('amenity') or negative ('disamenity'). Since our model handles all cases (i.e. production or consumption amenity; production or consumption disamenity), it has allowed us to design an identification procedure to figure out which case receives empirical support based on cross-city wage and rent variations. Moreover, by allowing for both firm and labor mobility, our model has also stressed the problem of endogeneity: does diversity causes wage and rent changes or vice-versa?

We have estimated the model through wage and rent regressions across US cities in the period 1970-1990. Either higher wage or higher rent or both have been shown to be significantly correlated with richer diversity. This result has survived several robustness checks against possible alternative explanations based on omitted variables.

To investigate the direction of causality of the above correlations, we have proposed to instrument diversity in two different ways, both exploiting the fact that, presumably exogenously from the characteristics of any single city, the overall migration into the US increased significantly between 1970 and 1990. Both instruments support the view that, during the period of observation. Richer diversity caused higher wages or higher rents or both for US-born residents across US cities.

Given our identification procedure, these findings are consistent only with a dominant production amenity effect of diversity: *a more multicultural urban environment makes US-born citizens more productive.* The choice of US cities as units of observation makes this result clean from most institutional differences that are generally shown to drive comparable cross-country studies. To the best of our knowledge, in terms of both data and identification procedure, our results are new. These results shed new light on the ongoing policy debate on the opportunity of imposing additional restrictions to migration flows in developed countries

It is worth mentioning, as a concluding remark, that while we established the positive effect of foreign-born residents, we did not "open the black box" to analyze theoretically and empirically what are the channels through which this effect works. We mentioned in the paper two interesting avenues through which this effect may work and we briefly qualify them here. First, foreign-born may have skills which are complementary to those of US-born. Even at the same level of education, problem solving, creativity and adaptability may differ between native and foreign-born workers so that reciprocal learning may take place. Second, foreign-born workers may provide services which are not perfectly substitutable with those of natives. An Italian stylist, a Mexican cook, a Russian dancer, simply provide different services than their US-born counterparts and, because of a taste for variety, this may increase the value of total production. The first effect described is likely to be stronger for the highly educated and in High Tech sectors. The second effect is likely to be important in the service sector. Suggestively, we run two separate specifications as (14) using one time the wage of low skilled (High School educated) and the other time that of high skilled (College educated) US-born white males 40 -50 as dependent variable[10]. The effect of the share of foreign-born is significantly positive on the high skilled returns (0.45 std. error 0.16) and zero on low-skilled

[10]The results of these last few regressions are not reported in the tables but are available on request.

returns (0.16 std. error 0.23). This result suggests that foreign-born's skills are more complementary to highly skilled US workers and explains also why most of the research that focused on low-skilled workers (Borjas 1999, Card 1990) did not find any positive effect of foreign-born. As suggestive exploration of the second channel, we run specification (14) separately on three subsamples. One only for the Low-Tech Manufacturing sector, another for the High-Tech Manufacturing sector and the third for the Service sector. Interestingly, the effect of the share or foreign-born workers on the average wage of US-born (white, male, 40-50) was 0.02 (std. error 0.19) in Low Tech Manufacturing, 0.55 (std. error 0.30) in High tech Manufacturing and 1.00 (std. error 0.19) in the service sector. This again is consistent with the idea that foreigners provide valuable complementary skills in High Tech manufacturing, but not in Low Tech manufacturing. Even more, they provide services which are not close substitutes with those of natives and increase substantially the productivity in the service sector. Overall we find our findings plausible and encouraging, leaving to future research the important goal of pursuing further the analysis of the mechanisms through which foreign-born residents affect the US economy.

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A Data Appendix

A.1 Data for MSA's

The data on cultural diversity and foreign-born are obtained from the 1970-1990 Public Use Microdata Sample (PUMS) of the US Census. We selected all people in working age (16-65) in each year and we identified the city where they lived using the SMSA code for 1990, while in 1970 we used the county group code to identify the metropolitan area. We used the variable 'Place of birth' in order to identify the country of origin of the person. We consider only the countries of origin that generate at least 0.5 percent of all foreign-born in the US working age population obtaining 35 groups for 1970 and 40 for 1990.

We use the Variable 'Salary and Wage' to measure the yearly wage income of each person. We transform the wage in real 1990 US $ terms by deflating it for the GDP deflator. To construct the average wage at the MSA level we select only white male US-born individual between 40 and 50 years of age. They constitute our group of homogeneous US-born citizen. The years of schooling for individuals of this group are measured using the variable 'higrad' for the 1970 census, which indicates the highest grade attended, while for 1990 the variable 'grade completed' is converted into years of schooling using Park (1994) correspondence Table 4. Average rents paid by the US-born are calculated using the group of white US-born males in working age and averaging by MSA their gross monthly rent per room (i.e. Rent divided by number of rooms) expressed in real 1990 US $ terms. The data on total city employment, total local public spending, and public spending in education are from the "County and City Databook".

The list of metropolitan areas used in our study is reported in the following table.

Name and state of the cities used

Abilene, TX	Dayton-Springfield, OH	Lexington, KY	Rockford, IL
Akron, OH	Decatur, IL	Lima, OH	Sacramento, CA
Albany-Schenectady-Troy, NY	Denver, CO	Lincoln, NE	Saginaw-Bay City-Midland, MI
Albuquerque, NM	Des Moines, IA	Little Rock-North Little Rock, AR	St. Louis, MO-IL
Allentown-Bethlehem-Easton, PA	Detroit, MI	Los Angeles-Long Beach, CA	Salem, OR
Altoona, PA	Duluth-Superior, MN-WI	Louisville, KY-IN	Salinas, CA
Amarillo, TX	El Paso, TX	Lubbock, TX	Salt Lake City-Ogden, UT
Appleton-Oshkosh-Neenah, WI	Erie, PA	Macon, GA	San Antonio, TX
Atlanta, GA	Eugene-Springfield, OR	Madison, WI	San Diego, CA
Atlantic-Cape May, NJ	Fayetteville, NC	Mansfield, OH	San Francisco, CA
Augusta-Aiken, GA-SC	Flint, MI	Memphis, TN-AR-MS	San Jose, CA
Austin-San Marcos, TX	Fort Lauderdale, FL	Miami, FL	Santa Barbara-Santa Maria-Lompoc, CA
Bakersfield, CA	Fort Wayne, IN	Milwaukee-Waukesha, WI	Santa Rosa, CA
Baltimore, MD	Fresno, CA	Minneapolis-St. Paul, MN-WI	Seattle-Bellevue-Everett, WA
Baton Rouge, LA	Gainesville, FL	Modesto, CA	Shreveport-Bossier City, LA
Beaumont-Port Arthur, TX	Gary, IN	Monroe, LA	South Bend, IN
Billings, MT	Grand Rapids-Muskegon-Holland, MI	Montgomery, AL	Spokane, WA
Biloxi-Gulfport-Pascagoula, MS	Green Bay, WI	Muncie, IN	Springfield, MO
Binghamton, NY	Greensboro--Winston-Salem--High Point, NC	Nashville, TN	Stockton-Lodi, CA
Birmingham, AL	Greenville-Spartanburg-Anderson, SC	New Orleans, LA	Syracuse, NY
Bloomington-Normal, IL	Hamilton-Middletown, OH	New York, NY	Tacoma, WA
Boise City, ID	Harrisburg-Lebanon-Carlisle, PA	Newark, NJ	Tampa-St. Petersburg-Clearwater, FL
Brownsville-Harlingen-San Benito, TX	Honolulu, HI	Norfolk-Virginia Beach-Newport News, VA-NC	Terre Haute, IN
Buffalo-Niagara Falls, NY	Houston, TX	Odessa-Midland, TX	Toledo, OH
Canton-Massillon, OH	Huntington-Ashland, WV-KY-OH	Oklahoma City, OK	Trenton, NJ
Cedar Rapids, IA	Indianapolis, IN	Omaha, NE-IA	Tucson, AZ
Champaign-Urbana, IL	Jackson, MI	Orlando, FL	Tulsa, OK
Charleston-North Charleston, SC	Jackson, MS	Pensacola, FL	Tuscaloosa, AL
Charlotte-Gastonia-Rock Hill, NC-SC	Jacksonville, FL	Peoria-Pekin, IL	Tyler, TX
Chattanooga, TN-GA	Jersey City, NJ	Philadelphia, PA-NJ	Utica-Rome, NY
Chicago, IL	Johnstown, PA	Phoenix-Mesa, AZ	Vallejo-Fairfield-Napa, CA
Cincinnati, OH-KY-IN	Kalamazoo-Battle Creek, MI	Pittsburgh, PA	Waco, TX
Cleveland-Lorain-Elyria, OH	Kansas City, MO-KS	Portland-Vancouver, OR-WA	Washington, DC-MD-VA-WV
Colorado Springs, CO	Kenosha, WI	Raleigh-Durham-Chapel Hill, NC	Waterloo-Cedar Falls, IA
Columbia, MO	Knoxville, TN	Reading, PA	West Palm Beach-Boca Raton, FL
Columbia, SC	Lafayette, LA	Reno, NV	Wichita, KS
Columbus, OH	Lafayette, IN	Richmond-Petersburg, VA	Wilmington-Newark, DE-MD
Corpus Christi, TX	Lancaster, PA	Riverside-San Bernardino, CA	Wilmington, NC
Dallas, TX	Lansing-East Lansing, MI	Roanoke, VA	York, PA
Davenport-Moline-Rock Island, IA-IL	Las Vegas, NV-AZ	Rochester, NY	Youngstown-Warren, OH

A.2 Grouping by Country of Birth

In Tables 1-9 we consider the diversity index constructed using 35 groups, corresponding to the 35 country of origin of immigrants which top the list of all countries of origin plus a residual group called 'others'. These 35 countries account for more than 90 percent of all foreign-born, both in 1970 and 1990, and a country that is not in this list supplies not more than 0.5 percent of all foreign-born living in the US. Here is the list of the non-residual countries, in alphabetical order. For year 1970 the countries are: Argentina, Australia, Canada, Czechoslovakia, China, Colombia, Cuba, Dominican Republic, England, France, Germany, Greece, Hungary, India, Ireland, Italy, Jamaica, Japan, Korea, Latvia, Lithuania, Mexico, Netherlands, Norway, Philippines, Poland, Portugal, Romania, Scotland, Sweden, Syria, Ukraine, URSS, Yugoslavia, Others. For 1990 the countries are: Argentina, Canada, China, Colombia, Cuba, Dominican Republic, Ecuador, England, France, Germany, Greece, Guyana, Haiti, Honduras, Hong-Kong, India, Iran, Ireland, Italy, Jamaica, Japan, Korea, Mexico, Nicaragua, Panama, Peru, Philippines, Poland, Portugal, El Salvador, Taiwan, Trinidad and Tobago, USSR, Vietnam, Yugoslavia.

In Tables 10 and 11, to have the same group classification in 1970 and 1990, we allocate more than one non-residual country to the same group based on geographical proximity. Our fifteen groups are almost the same as those defined and used in Card (2001), Table 5. This is the list: Mexico, Caribbean Countries, Central America, China-Honk Kong-Singapore, South America, South East Asia, Korea and Japan, Philippines, Australia-New Zealand-Canada-UK, India and Pakistan, Russia and Central Europe, Turkey, North Africa and Middle East, Northwestern Europe and Israel, South-western Europe, Sub-Saharan Africa, Cuba.

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