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The economic value of cultural diversity: evidence from US cities

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Abstract

What are the economic consequences to U.S. natives of the growing diversity of American cities? Is their productivity or utility affected by cultural diversity as measured by diversity of countries of birth of U.S. residents? We document in this paper a very robust correlation: US-born citizens living in metropolitan areas where the share of foreign-born increased between 1970 and 1990, experienced a significant increase in their wage and in the rental price of their housing. Such finding is economically significant and survives omitted variable bias and endogeneity bias. As people and firms are mobile across cities in the long run we argue that, in equilibrium, these correlations are consistent with a net positive effect of cultural diversity on the productivity of natives.

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1. Introduction

Since the 1965 amendments to the Immigration and Nationality Act immigration into the United States has been on an upward surge. Indeed, immigration rates have been accelerating since the eighties. As a consequence, during the last thirty years foreign born residents in the United States have increased substantially as a share of both the total population and the labor force. In 1970 only 4.8% of the US residents were foreign-born; that percentage grew to 8% in 1990 and to 12.5% in the year 2000. Similarly, although to a lesser extent, other industrialized countries such as Europe and Australia have also recently experienced rising pressures from immigrants.[1] This phenomenon has spurred a heated policy debate and galvanized academic interest.

There is a large and growing body of empirical literature on the consequences of migration (see, among others Borjas 1994, 1995, 1999, 2003; Borjas et al., 1997; Boeri et al., 2002; Card 1990, 2001; Card and Di Nardo, 2000). This literature, however, has disproportionately focussed on one aspect of the subject: the impact of low-skilled immigrants on US wages. These studies typically treat labor markets for different skills as segmented, and focus on the consequences of wages for different skill-groups in the

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1 See Peri (2005) for a comparison of immigration in the US and in the EU during the nineties.

short and medium run. Our work takes a different angle. Rather than study the short-run effects of new immigrants on the receiving country in a classic model of skill supply and demand, we consider a simple multi-city model of production and consumption in order to ask 'what is the economic value of "diversity" that the foreign born bring to each city'. The foreign born conceivably have different sets of skills and abilities than the US born, and therefore could serve as valuable factors in the production of differentiated goods and services. As different US cities attract very different shares of foreign-born we can learn about the value of such 'diversity' from the long-run equilibrium distribution of wages and prices across cities. For the rest of the paper, the term 'cultural diversity' will refer to the diversity of the workers' countries of birth (rather than ethnicity or ancestry characteristics) and will be measured by an index of 'plurality' of countries of origin.

Diversity over several dimensions has been considered by economists as valuable both in consumption and production. Jacobs (1969) attributes the prosperity of cities to their industrial diversity. Quigley (1998) and Glaeser et al. (2001) identify the diversity of available consumption goods and services as one of the attractive features of cities. Florida (2002a, 2002b) stresses the importance of the diversity of creative professions employed in research and development or high tech industries. More generally, Fujita et al. (1999) use the 'love of variety' in preferences and technology as the building block of their theory of spatial development: the production of a larger variety of goods and services in a particular location increases the productivity and utility of people living in that location.

Against this background, we conjecture that cultural diversity may very well be an important aspect of urban diversity, influencing local production and/or consumption.[2] The aim of this paper is to test this conjecture by quantifying the value of cultural diversity to US-born people. Who can deny that Italian restaurants, French beauty shops, German breweries, Belgian chocolate stores, Russian ballets, Chinese markets, and Indian tea houses all constitute valuable consumption amenities that would be inaccessible to Americans were it 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 in the workplace.[3] Cultural diversity, therefore, may increase consumption variety and improve the productivity of natives. On the other hand, natives may not enjoy living in a multi-cultural environment if they feel that their own cultural values are being endangered. Moreover, intercultural frictions may reduce productivity, particularly if natives associate increasing immigration with further job losses for the US born. Thus cultural diversity could possibly decrease both the utility and the productivity of natives.

We focus on 160 major metropolitan areas in the US, for which we can construct consistent data between 1970 and 1990. While these metropolitan areas do not cover

2 An economically oriented survey of the pros and cons of ethnic diversity is presented by Alesina and La Ferrara (2003).

3 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% (14 laureates) were not US-born. During the same time period the share of foreign-born in the general population was on average only 10%. From our perspective, such example is interesting because research in hard sciences is typically based on large team work.

the whole US urban population, they include the largest and most important cities. More importantly, they span the whole range of 'diversity', for they include the most diverse cities (New York, Los Angeles, San Francisco) along with some of the least diverse. We use the 'index of fractionalization' (by the country of birth of each city resident) in order to measure cultural diversity across these 160 cities.[4] This index measures the probability that, in any one city, two individuals chosen at random were born in different countries. Cities entirely populated by US-born individuals would have an index of fractionalization equal to 0. Going to the other extreme, if each individual within a city was born in a different country, the index would equal one. US cities vary wildly by this measure, ranging from 0.02 (Cleveland) to 0.58 (Los Angeles). Since US-born people are highly mobile across US cities, following Roback (1982) we develop a model of 'open cities' that allows us to use the observed variations of wages and rents of US-born workers to identify the production and consumption gains associated with cultural diversity. In particular, we estimate two regressions in which cultural diversity, measured as 'fractionalization' (or the share of foreign-born residents) affects the average wage received and the average rent paid by US-born workers. Our main finding is that, on average, *cultural diversity has a net positive effect on the productivity of US-born citizens* because it is positively correlated with both the average wage received and the average rent paid by US-born individuals. This partial correlation survives the inclusion of many variables that proxy for productivity and amenity shocks across cities.

Two fundamental concerns arise when we attempt to interpret these correlations as *causal* effects of diversity on the wages and rents of natives, namely a potential endogeneity bias and the possibility of spatial selection of natives. Endogeneity works as follows. Cities may experience an increase in the average wage from a positive economic shock, disproportionately attracting immigrants and thus witnessing an increase in diversity (this hypothesis is often referred to as 'boom cities'). If this were the true story, the measured impact of diversity on wages and rents would be upwardly biased. To tackle this problem, we use instrumental variable estimations, a method widely used among economists that requires an 'auxiliary' variable whose exogenous variation affects diversity in a city (but not its productivity). Such a variable allows us to isolate that portion of the correlation between diversity and wages that is due to the causal effect of diversity on wages.

The spatial selection of native workers, on the other hand, is harder to deal with. In fact, if the presence of foreign-born people attracts a particular type of US born worker (call this group 'tolerant') and these workers also happen to be more productive, then the correlation between diversity and productivity of natives may be the effect of this selection rather than of complementarities or externalities with foreign-born. The best we can do is to control for observable characteristics of US-born residents and assume that their 'tolerance' is not highly correlated with the residual (unobserved) productivity. This issue, however, is certainly not settled with this paper and needs more research. We will come back to it in the final part of the paper.

The rest of the paper is organized as follows. Section 2 reviews the literature on the economic consequences of immigration and cultural diversity. In particular we

4 As an alternative and perhaps more intuitive measure of diversity in a city we also use, in several parts of the analysis, the share of foreign-born residents.

differentiate our work from (and reconcile it with) the common findings in labor economics that immigrants have negative or zero effects on the wages of US-born workers. Section 3 introduces our dataset and surveys the main stylized facts. Section 4 develops the theoretical model that is used to design and interpret our estimation strategy. Section 5 presents the results from the basic estimation, checks their robustness and tackles the issue of endogeneity. Section 6 discusses the results and provides some important caveats and qualifications to our conclusions.

2. Literature on diversity

Cultural diversity is a broad concept that has attracted the attention of both economists and social scientists. The applied 'labor' literature has analyzed ethnic diversity and ethnic 'segregation' in the US, as well as their impact on economic discrimination and the achievements of minorities.[5] The present paper does not focus on this aspect of cultural diversity even though we control for black-white composition issues.

More closely related to our analysis is the literature concerning the impact of immigration on the US labor market. Several contributions by George Borjas (notably Borjas, 1994, 1995, 1999, 2001 and 2003) focus on the issue of US immigration as a whole, and its effect on native workers. Similarly, important contributions by David Card (notably, Card, 1990; Butcher and Card, 1991; Card and Di Nardo, 2000; Card, 2001) analyze the wages and reactions of domestic workers to inflows of new immigrants by exploiting the geographic variation of immigration rates and wages across US states or US cities. These contributions do not achieve a consensus view either on the effect of new immigrants on the wages of domestic workers (which seems small except, possibly, for low skill levels) or on the effect of new immigrants on the migration behavior of domestic workers. Let us emphasize, however, that the negative (significant or small) effect that is found in this literature is merely a 'relative' effect. Immigrants bring down the relative wages of low-skilled workers (but raise the wages of intermediately-skilled workers) due to their composition (abundant in low skills and scarce in intermediate skills). This, however, does not comment on the overall (average) effect on US workers. In the presence of complementarities between the skills of immigrants and the skills of natives, or of externalities from highly skilled workers (who are also abundant among immigrants), the impact of immigration on the average wage of US born workers may very well be positive. While the labor literature estimates the relative effect of immigration within labor markets segmented by skills (such an effect would be negative if different skills are imperfect substitutes), we focus on the average effect of immigration that results from aggregating those effects with the positive complementarity-effects and the positive externality-effects.[6] This is a novel approach, and while we do not deny that a shift of relative wages (between skills) takes place as a consequence of immigration, we focus on the average overall effect on wages of US-born workers

5 Notable examples are Card and Krueger (1992, 1993), Cutler and Glaeser (1997), Eckstein and Wolpin (1999), Mason (2000).

6 While in the present paper we simplify these effects into an overall effect of diversity on the TFP of US-born workers, in Ottaviano and Peri (2005) we separately model and analyze the effects of complementarieties across skills. We find that the (positive) empirical effects of migration on the average wage of US-born workers are very close to the theoretically calculated effects from the diversity of skills generated by immigrants.

and find it significantly positive. Recently, evidence of a positive effect of immigrant inflows on rents in cities has been provided by Saiz (2003a, 2003b), although he interprets this as a consequence of increased demand in housing rather than an increased value of houses due to higher diversity and higher wages. To our knowledge this is the first work that looks at a general equilibrium effect of immigration (diversity) on wages, employment and rents of US born residents.

In short, the standard labor literature assumes that immigrants and domestic workers within a particular skill group are homogeneous, so that immigration will shift the labor supply and change local wages in that skill group, the extent of which will depend on the mobility of domestic workers. Our approach takes a rather different stand. We believe that 'place of birth' can be a feature that differentiates individuals in terms of their attributes, and that this differentiation may have positive or negative effects on the productivity (through complementarities and externalities) and the utility (through taste for variety) of US-born residents. Moreover, we consider equilibrium variations of wages and rents in the long-run, relying on the assumption of mobility of native workers and firms across cities.

Relevant to our work, several researchers in the social sciences have related diversity with urban agglomeration. The functioning and thriving of urban clusters relies on the variety of people, factors, goods and services within them. Examples abound in the urban studies literature. Jacobs (1969) views 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. A key feature of these cities is the cultural diversity of their populations. Similarly, Bairoch (1988) sees cities and their diversity as the engines of economic growth. Such diversity, however, has been seen mainly in terms of the diversified provision of consumer goods and services, as well as productive inputs (see, e.g. Quigley, 1998; Glaeser et al., 2001). In his work within the nexus of sociology and economics, Richard Florida (2002a, 2002b) argues that 'diverse' and tolerant cities are more likely to be populated by creative people, thus attracting industries such as high tech and research that heavily rely on creativity and innovative ability. The positive 'production value' of diversity has also been stressed in the literature on the organization and management of teams. Here the standard assumption is that higher diversity can lead to more innovation and creativity by increasing the number of ways groups frame problems, thus producing a richer set of alternative solutions and consequently better 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 whose countries of origin have different cultures, legal systems, and languages imposes costs on the firm that would not be present if all the workers had similar backgrounds. However, complementarity between workers, in terms of skills, can more than offset the costs of cross-cultural interaction.[7]

Finally, several studies in political economics have looked at the historical effects of cultural and ethnic diversity on the formation and quality of institutions.

7 Berliant and Fujita (2004) 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. In this respect, a perpetual reallocation of members across different teams may be necessary to keep creativity alive.

The traditional wisdom (confirmed by Easterly and Levine, 1997) had been that more fragmented (i.e. diverse) societies promote more conflicts and predatory behavior, stifling economic growth. However, recent studies have questioned that logic by showing that higher ethnic diversity is not necessarily harmful to economic development (see, e.g., Lian and Oneal, 1997). Collier (2001) finds that, as long as institutions are democratic, fractionalized societies perform better in the private sector than more homogenous ones. Framed within efficient institutions, diversity may serve as a valuable asset for society.

3. Cultural diversity, wages and rents

The questions we are interested in are the following. How does cultural diversity affect the US-born? Do they benefit or loose from the presence of foreign-born? How do we measure such benefits or costs?

We are able to extract interesting insights into these questions by analyzing the wage and rent distributions across cities, assuming that such distributions are the equilibrium outcomes of economically motivated choices. We assume that workers and firms are mobile across cities, and so can change their location in the long run if a productivity shock or a price differential were to arise. Since people can 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 as they have eliminated any systematic difference in indirect utility and profits through migration. Before formalizing these ideas in Section 4, we put our theoretical analysis into context by introducing our measure of cultural diversity (Section 3.1) and by establishing the main stylized facts about wages, rents and diversity in US cities (Section 3.3).

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 the years 1970 and 1990 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 and spending on local public goods. We consider 160 Standard MSA's that could be consistently identified in each census year. Our dataset contains around 1,200,000 individual observations for 1990, and 500,000 for 1970. We use these to construct aggregate variables and indices at the MSA level. The reasons for focusing on metropolitan areas are two-fold. First, urban areas constitute closely connected economic units within which interactions are intense. Second, they exhibit a higher degree of diversity than non-urban areas because immigrants traditionally settle in large cities. While it is possible to construct data only on 160 metropolitan areas (using 1970 and 1990 PUMS of the US Bureau of Census) those areas include the most important US cities, spanning a wide range of variation in terms of cultural diversity. Adding all the other metropolitan areas would simply amount to adding more observations characterized by low and similar levels of diversity. This would certainly add some noise, but probably would not help much in the identification of the effect of diversity on wages and rents.

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. We denote by $\bar{w}_{US,c,t}$ the resulting average wage for city c in year t. This value is neither affected by composition effects nor distorted by potential discrimination factors (across genders or ethnicity) or life-cycle considerations. It can therefore serve as a good proxy for the average wage of US-born workers in the city, comparable across census years. The correlation between $\bar{w}_{US,c,t}$ and the degree of diversity of a city comes only through the equilibrium effect of diversity on the labor demand and supply of native workers. As a measure of the average land rent in an MSA we use the average monthly rent paid per room (i.e. the monthly rent divided by the number of rooms) by white US-born male residents of working age (16–65 year).[8] We denote this measure (for city c in year t) as $\bar{r}_{US,c,t}$. While this measure does not control for housing quality (beyond the number of rooms), there is no reason to think that housing quality is related to the percentage of foreign-born in a city, so this measure should not induce any relevant bias in the relation.

Turning to our key explanatory variable, our measure of cultural diversity considers the country of birth of people as defining their cultural identity. Foreign born residents have always been an important part of the US population, and their share of the population has only grown larger in the past decades. In 1970, they constituted 4.8% of the total population, while in 1990 they reached 8%, still continuing to grow afterwards. Our measure of cultural diversity is the so called 'index of fractionalization' (henceforth, simply 'diversity index'), routinely used in the political economics literature. This index has been popularized by cross-country studies by Mauro (1995) and has been widely used since. The index is simply 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).[9] Specifically, we use the variable CoB (Country of Birth of a person) to define the cultural identity of each group. The diversity index is defined as:

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

where $\left(CoB_i^c\right)_t$ is the share of people born in country i among the residents of city c in year t. This index is an increasing measure of both the cultural 'richness' of a city (i.e. the number of groups) and its cultural 'diversity' (i.e. the evenness of groups' sizes). It reaches its minimum value 0 when all individuals are born in the same country, and its maximum value 1 when there are no individuals born in the same country. Intuitively, when all individuals belong to the same group, the probability that two randomly selected individuals belong to different groups is 0, whereas it equals 1 when all individuals belong to

8 The housing market is less segmented by skills than the labor market. Therefore we use a larger age-range in order to calculate average rents.

9 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.

Table 1. Foreign Born living in 160 U.S. metropolitan areas 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.

different groups. On the other hand, for a given number of groups M (i.e. controlling for 'richness'), the index reaches its maximum at $(1 - 1/M)$ when individuals are uniformly distributed across groups.[10]

The 1970 and 1990 PUMS data report the country of birth of each individual. We count as separate groups the migrants of each country of origin contributing at least 0.5% of the total foreign-born population working in the US. Migrants from other countries of origin are gathered in a residual group. This choice implies that we consider 35 countries of origin both in 1970 and in 1990. These groups constitute 92% of all foreign-born immigrants; the remaining 8% are merged into a single group. The complete list of countries for each census year is reported in the data appendix, while the largest 15 of these groups are reported in Table 1. As the Table shows, between 1970 and 1990, the origin of immigrants has increasingly become Mexico; the share of foreign born, however, has increased as well, so that overall the diversity index has increased. As to the main countries of origin of immigrants, we note the well known shift from European countries towards Asian and Latin American countries.

3.2. Diversity across US cities

Table 2 shows the percentage of foreign-born and the diversity index for a representative group of metropolitan areas in the year 1990. To put into context the extent of

10 In our case as M, the number of groups, is 36 the maximum for the index is 0.972. See Maignan et al. (2003) for further details.

Table 2. Diversity in representative Metropolitan Areas, 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 1990 PUMS census data.

diversity across US cities, each diversity index can be compared with the cross-country value of the index of linguistic fractionalization reported by the Atlas Narodov Mira and published in Taylor and Hudson (1972) for the year 1960. These values have been largely used in the growth literature (see e.g. Easterly and Levine, 1997; Collier, 2001). Since foreign-born immigrants typically use their country's mother tongue at home, thus signalling their country's cultural identity, our diversity index captures cultural and linguistic fragmentation for different US cities much as that index does for different countries in the world. The comparison is instructive. Diversified cities, such as New York or Los Angeles, have diversity indices between 0.5 and 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. More homogenous cities, such as Cincinnati and Pittsburgh, exhibit a degree of fractionalization of only 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 the nations of the world.

Table 2 also shows that, even though people born in Mexico constitute an important group in many cities, the variety of countries of origin of residents of US cities is still remarkable. Finally we note that there is a very high correlation between the diversity index and the share of foreign born in a city. The main reason an American city is considered 'diverse' is because there is a large percentage of foreign born living there, not necessarily because there is a high degree of diversity within the foreign born.

3.3. Stylized facts

The key empirical finding of our paper is readily stated: *ceteris paribus, US-born workers living in cities with higher cultural diversity are paid, on average, higher wages, and pay higher rents, than those living in cities with lower cultural diversity*. In Section 5 we show that this correlation not only survives the inclusion of several other control variables, but it is likely to be the result of causation running from diversity to wages and rents.

We report in Figures 1 and 2, below, the correlation between the change of the diversity index for the 1970–1990 period, $\Delta(div_{c,t})$, and the percentage change in the wage of the US-born, $\Delta\ \ln\left(\bar{w}_{US,c}\right)$, or the percentage change in rents paid by the US-born, $\Delta\ \ln\left(\bar{r}_{US,c}\right)$ in 160 metropolitan areas. The effect of fixed city characteristics, such as location or geographic amenities, are eliminated by differencing. The figures show the scatter-plots of these partial correlations and report the OLS regression line. Cities whose diversity increased more than the average, during the 20 years considered (such as Jersey City, Los Angeles, San Francisco, and San Jose), have also experienced larger than average wage increases for their US-born residents. Similarly they also experienced a larger than average increase in rents. The OLS coefficient estimates imply that a city experiencing an increase of 0.09 in the diversity index (as Los Angeles did) would experience associated increases of 11 percentage points in the average wage and 17.7 percentage points in the average rent paid by US-born residents, relative to a city whose diversity index did not change at all (such as Cleveland).

4. Theoretical framework

4.1. The model

To structure and interpret our empirical investigation, we develop a stylized model in which 'diversity' affects both the productivity of firms and the satisfaction of consumers through a localized effect. Both the model and the identification procedure build on Roback (1982).[11]

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. We assume that inter-city commuting costs are prohibitive, so that for all workers the city of work and residence coincides. We also ignore intra-city commuting costs, which allows us to focus on the inter-city allocation of workers.

The overall amount of labor available in the economy is equal to L. It is inelastically supplied by urban residents; 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 who work and reside in city c. Workers are all identical in terms of attributes that are

11 Roback's (1982) framework has been extensively applied to measure the value of local amenities or local factors of production. Examples include Rauch (1993), Kahn (1995), and Dekle and Eaton (1999).

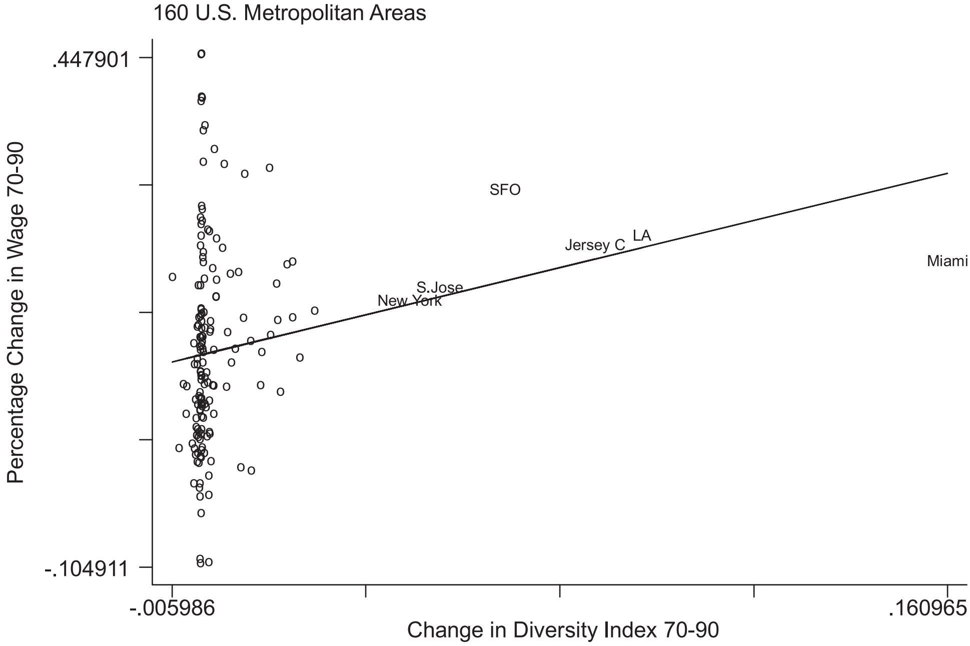

Figure 1. Wages of US-born and diversity.

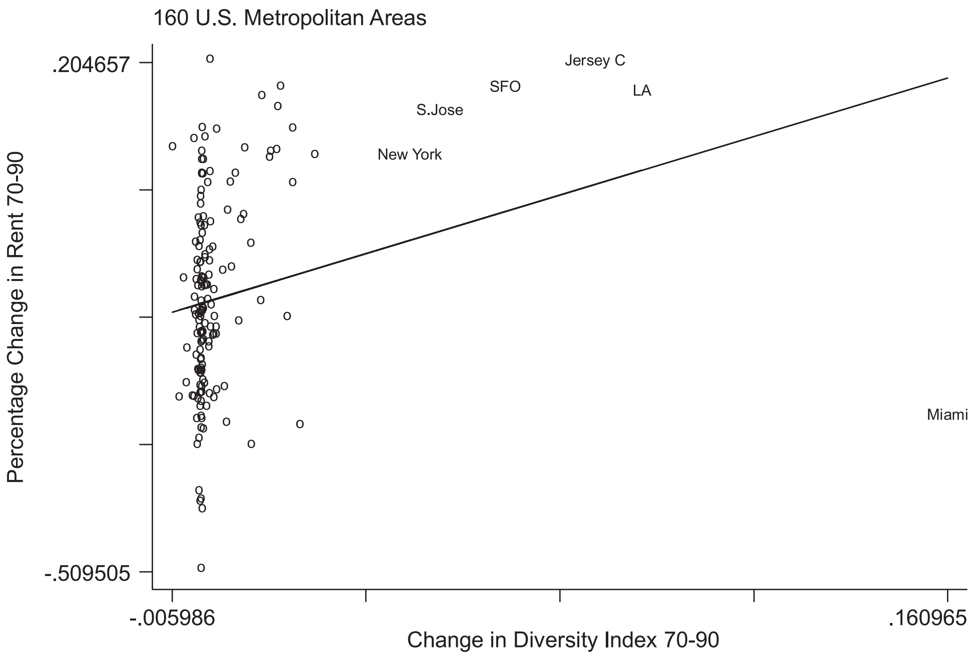

Figure 2. Rents of US-born and diversity.

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 size L_{ic} ('evenness') of resident groups, enters both production and consumption as an effect that, in principle, can be positive or negative. To establish the existence and the sign of such effect is the final aim of the paper. While land is fixed among cities, it is nonetheless mobile between uses within the same city.[12] 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.[13]

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 equation (2) H_{ic} and Y_{ic} are land and good consumption respectively, while $A_u(d_c)$ captures the 'utility effect' associated with local diversity d_c. If the first derivative $A_u'(d_c)$ is positive, diversity can be seen as a local amenity; if negative as a local dis-amenity.

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

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

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 equation (5) H_{jc} and L_{jc} are land and labor inputs respectively. $A_Y(d_c)$ captures the 'productivity effect' associated with local diversity d_c. It is convenient to treat the effect of diversity as a shift in total factor productivity, $A_Y'(d_c)$, that is

12 The assumption of exogenous and constant land area of a city is harmless. The same implications would follow under the more realistic assumption that expanding the land area of a city comes at a cost because of internal commuting costs and lower quality of the marginal land.

13 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 location, and thus does not affect migration choice. The alternative assumptions of absentee landlords or balanced ownership of land across all cities would also serve that purpose.

common to all firms in city c. This shift could be positive or negative.[14] We should notice at this point that assuming identical effects of diversity on utility, $A_U(d_c)$, and productivity, $A_Y(d_c)$, across agents (i) and firms (j) is critical in order to use the model by Roback (1982) to characterize the average equilibrium rent and wage as a function of only diversity. If diversity were to affect firms and agents in different ways (say because some people like diversity more than others and some firms need diversity of workers more than others) then in equilibrium US-born agents would sort themselves across cities (see e.g. Combes et al., 2004). In this case the equilibrium wages and rents across cities would reflect not only different levels of diversity but also different evaluations of diversity by US-born individuals and firms. Such an equilibrium with heterogeneous agents would complicate the use of average wages and rents to infer the impact of diversity on productivity. The analysis of diversity assuming heterogeneous effects on US born agents is certainly an interesting issue that we leave for future research.

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

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

which implies marginal cost pricing:

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

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: $E_{ic} = \omega_c$. Since good Y is freely traded, its price is the same everywhere. We choose this good as numeraire, which allow us to write p_c=1.[15]

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

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

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

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

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

14 The contribution of diversity to total factor productivity could stem from imperfect substitutability of different groups as well as from pecuniary or learning externalities. For instance, Ottaviano and Peri (2004a) derive a production function similar to equation (5) with non-tradable intermediates and taste for variety.

15 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 is supported by the large positive correlation between local price indices and land rents at the SMSA level.

To complete the equilibrium analysis 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 $Y_c = \sum_j Y_{jc}$, equation (6) implies $\omega_c L_c = \alpha p_c Y_c$. Given $H_c = \sum_j H_{jc} + \sum_i H_{ic}$, equation (6) and (3) imply $\mu r_c H_c = (1-\alpha \mu) p_c Y_c$. Together with $E_{ic} = \omega_c$ and $p_c = 1$, these results can be plugged into equation (4) to obtain:

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

Equation (10) shows that the indirect utility of a person is higher, *ceteris paribus*, in a city with low population density, L_c/H_c, (because of the lower price of housing) and is affected by diversity through its impact on productivity, $A_Y(d_c)$, which determines wages, and its direct effect on utility $A_U(d_c)$. Substituting equation (10) into equation (9) generates a system of equations that can be solved for the equilibrium spatial allocation of workers. In particular, 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 $i = 1, \ldots, M$ to each city $c = 1, \ldots, N$. Constant returns to scale and fixed land ensure that the spatial equilibrium is unique and has a positive number of workers in every city ('no ghost town'). Then, the composition of the urban community depends on the net impact of diversity on utility and productivity.

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 as in equation (8) and the free migration conditions in equation (9) that take equation (4) into account. Specifically, call v the equilibrium value of indirect utility. Due to the free mobility of US-born individuals, this value is common among cities and, due to the large number of cities, is unaffected by city-level idiosyncratic shocks. Then, solving equations (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 foundation of our empirical analysis. They capture the equilibrium relationship between diversity and factor prices. In light of Roback (1982), the two equations must be estimated together in order to identify the effects of diversity on productivity and utility. Consider, for instance, equation (11) in isolation. A positive correlation between d_c and r_c is consistent either with a positive effect of diversity on utility ($A_U'(d_c) > 0$) or a positive effect of diversity on productivity ($A_Y'(d_c) > 0$). Analogously, if one considers equation (12) in isolation, a positive correlation between d_c and w_c is consistent either with a negative utility effect ($A_U'(d_c) < 0$) or a positive productivity effect ($A_Y'(d_c) > 0$) from diversity. Only the joint estimation of equations (11) and (12) allows one to establish which effect indeed dominates. Specifically:

$$\begin{aligned}
&\frac{\partial r_c}{\partial d_c} > 0 \text{ and } \frac{\partial w_c}{\partial d_c} > 0 \text{ iff } \textit{dominant positive productivity effect } (A_Y'(d_c) > 0)\\
&\frac{\partial r_c}{\partial d_c} > 0 \text{ and } \frac{\partial w_c}{\partial d_c} < 0 \text{ iff } \textit{dominant positive utility effect } (A_U'(d_c) > 0)\\
&\frac{\partial r_c}{\partial d_c} < 0 \text{ and } \frac{\partial w_c}{\partial d_c} < 0 \text{ iff } \textit{dominant negative productivity effect } (A_Y'(d_c) < 0)\\
&\frac{\partial r_c}{\partial d_c} < 0 \text{ and } \frac{\partial w_c}{\partial d_c} > 0 \text{ iff } \textit{dominant negative utility effect } (A_U'(d_c) < 0)
\end{aligned} \tag{13}$$

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. Given the utility level v and diversity d_c, the free entry condition in equation (8) is met along the downward sloping curve, while the free migration condition in equation (9) holds along the upward sloping curve. The equilibrium factor prices for city c 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 equation (8) up (down) if diversity has a positive (negative) productivity effect and it shifts equation (9) up (down) if diversity has a positive (negative) utility 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

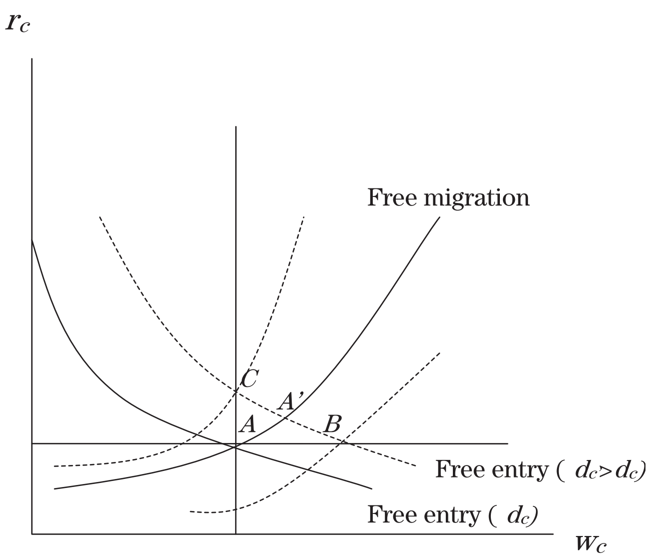

Figure 3. The spatial equilibrium.

upward shift of equation (8) dwarfs any shift of equation (9); i.e. the positive productivity effect dominates.

5. Wage and rent regressions

5.1. Basic specifications

The theoretical model above provides us with a consistent framework to structure our empirical analysis. In particular it suggests how to use wage and rent regressions to identify the effects of diversity, a characteristic particular to each city, on the productivity and utility of US natives. Our units of observation are the 160 Metropolitan Statistical Areas (MSA's) 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:

$$\ln(\bar{w}_{US,c,t}) = \beta_1(Controls_{c,t}) + \beta_2(div_{c,t}) + e_c + e_t + e_{c,t} \tag{14}$$

The average wage of natives in city c in year t, $\bar{w}_{US,c,t}$, is defined as described in Section 3.1. The focal independent variable is $div_{c,t}$, which is the diversity index defined in equation (1). The other independent variable, $Controls_{c,t}$, capture other controls. Specifically we always include among the controls some measure of the average education of workers in city c at time t (either the average schooling or the share of education groups) while in Section 5.2 we include several other alternative variables which may potentially affect the productivity and the share of foreign-born in a city. We also include 160 city fixed effects e_c and common time-effects e_t. Finally, $e_{c,t}$ is a zero-mean random error term independent from the other regressors.

Under these assumptions, the coefficient β_2 captures the equilibrium effect of a change in cultural diversity on wages. However, as discussed in subsection 4.2, the sign of β_2 cannot be directly interpreted as evidence of any positive effect of diversity on production. Identification thus requires us to estimate the following parallel rent regression:

$$\ln(\bar{r}_{US,c,t}) = \gamma_1(Controls_{c,t}) + \gamma_2 div_{c,t} + \varepsilon_c + \varepsilon_t + \varepsilon_{ct} \tag{15}$$

Our definition of the average rent per room of natives $\bar{r}_{US,c,t}$ in city c in year t is described in Section 3.1. The focal independent variable is again the diversity index $div_{c,t}$. The other independent variables, $Controls_{c,t}$, capture other controls. We add these to check that the correlation of interest is robust to the inclusion of other variables, and thus is not spurious. Further we control for city fixed effects ε_c, include a year dummy ε_t, and assume that $\varepsilon_{c,t}$ is a zero-mean random error uncorrelated with the regressors. The coefficient γ_2 captures the equilibrium effect of a change in cultural diversity on average city rents. By merging the information on the signs of β_2 and γ_2, we are able to identify the net effect of diversity. We begin by estimating the two basic regressions using least squares, including further controls and using different estimation methods later on as we proceed.

The least squares estimates of the regressions (14) and (15) are reported in specifications I and VII of Table 3. Specification I shows the basic estimates for the wage equation, when we only include, besides state and year fixed effects, the average schooling of the considered group of white US-born males 40–50 years of age as a control. Specification VII considers the rent equation with only state and year fixed effects as controls. The estimated coefficients β_2 and γ_2 are both positive and statistically and

Table 3. Basic Wage and Rent Specifications

Dependent variable	Average log wage for US-born workers						Average log rent for US-born residents		
Specification:	I Base 1 wage	II 4 school groups	III Polynomial school	IV Base 1, Pop. weighted	V Include empl.	VI Base 2 wage	VII Base 1 rent	VIII With population and income	XI Base 2 rent
Average schooling	0.11**			0.11*	0.11**	0.10**			
	(0.01)			(0.01)	(0.01)	(0.01)			
4 School groups		Yes							
Quartic in schooling			Yes						
ln(income per capita)								0.67**	
								(0.08)	
ln(employment)					0.02				
					(0.02)				
ln(population)								0.03	
								(0.04)	
Diversity index	1.27**	1.17**	1.29**	1.37**	1.29**		1.90*	0.95**	
	(0.30)	(0.36)	(0.30)	(0.23)	(0.29)		(0.60)	(0.50)	
Share of foreign born						0.57**			1.13**
						(0.11)			(0.24)
Diversity index among foreign born						0.14*			0.12
						(0.08)			(0.16)
City fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2 (excluding city and time fixed effects)	0.10	0.14	0.12	0.11	0.10	0.12	0.30	0.30	0.31
Observations	320	320	320	320	320	320	320	320	320

Specification I–VI: Dependent variable is logged average yearly wage of white, US-born, males 40–50 years expressed in 1990 US$.
Specification VII–IX: Dependent variable is logged average monthly rent per room paid by white, US born 16–65 years of age, expressed in 1990 US$.
**Significant at 5%, * significant at 10%.
In parenthesis: heteroskedasticity-robust standard errors.

economically significant. An increase in the diversity index by 0.1 (roughly the increase experienced by Los Angeles during the 1970–1990 period) is associated with a 13% increase in the average real wage of US natives and with a 19% increase in real rents.

Similarly specifications VI and IX of Table 3 use the same controls as specification I and VII and decompose the effect of diversity (on wages and rents) into two parts. Specifically the diversity index can be expressed as the contribution of two factors. First, a city is more diverse if the overall group of foreign-born people is larger. Second, a city is more diverse if the foreign-born group is made up of a wider variety of groups. The diversity index can thus be written as a (non-linear) function of the share of foreign-born, or a diversity index can be calculated considering only the foreign born. We enter these two factors separately in specifications VI and IX in order to analyze their impact on wages and rents, respectively. Let us note that the share of foreign born is, by far, the most important component in determining the variation of the diversity index across cities. It explains, by itself, almost 90% of the index variation. It is not a surprise, therefore, to find that the share of foreigners is the most important contributor to the effect on wages and rents. An increase in the share of foreign born by 0.25 (experienced by Los Angeles during the considered period) is associated with a 14.5% increase in wages of US natives and a 28% increase in rents. The effect of the diversity of foreigners, on the other hand, has a positive but hardly significant impact.

The intermediate specifications (II to V for the wage equations and VIII for the rent equation) in Table 3 include alternative controls in order to check wether the correlation is robust to potential omitted variables. Specification II of the wage regression controls for the schooling of the group of US-born by including the shares of three groups (high school graduates, college dropouts and college graduates) among the total employed in each city, rather than simply the average years of schooling. Specification III includes a quartic polynomial in average schooling. While non-linear effects at different schooling levels may be relevant, here we see that the coefficient on diversity changes only marginally when we use different methods to control for education. We also run a specification (not reported) controlling for individual years of schooling in the construction of $\ln\left(\bar{w}_{US,c,t}\right)$, rather than at the second stage. Doing this reduces the coefficient on diversity somewhat to 1.00 (standard error equal to 0.32). All in all how we control for education does not seem to have a relevant effect on the coefficient on diversity. Specification IV weighs each observation (city) by its population. This control allows us to under-emphasize the role of small cities. The effect of diversity does not change much with this amendment; in fact it increases a bit (the coefficient is now equal to 1.37), which is a consequence of the fact that cities in which diversity has the largest impact (as seen in Figure 1 and 2) are indeed the largest cities, such as Los Angeles and New York. Specification V includes the log of employment as an additional control. On the one hand, if there are effects of employment density on productivity (as suggested by Ciccone and Hall, 1996) it may be relevant to control for employment; on the other hand employment (along with wages and rents) is determined endogenously as an equilibrium outcome in our model. As a consequence, including an endogenous variable as a control may bias the estimates of all coefficients. Fortunately we find that employment is not significantly correlated to wages (coefficient equal to 0.02 with standard error equal to 0.02), and its inclusion does not change the coefficient on diversity much. Omitting employment, therefore, is theoretically justified and empirically sound. These specifications reassure us that our basic specification captures both the correct sign and magnitude of the correlation between diversity and wages.

As for the rent regression, column VIII includes the average log income and log population of each city as controls. In reality, these variables may depend on several exogenous factors and may affect the value of housing. They are, however, endogenously determined in the equilibrium described in Section (4). In fact wages are the main determinant of income, while population is affected by internal migration. The two channels through which diversity can affect rents, described by our model, are either by increasing productivity (which pushes up income and rents), or by increasing the desirability of a city. When controlling for income and population, a residual positive effect of diversity would imply that people do value diversity per se, and are willing to bid up rents more than what would be implied only by higher income and higher population. The problem, however, is that including these two endogenous variables may induce a bias in the estimates of the coefficients of regression in equation (15). The estimated coefficients in specification VIII show that including income and population reduces the effect of diversity by half. In particular income per capita is a main determinant of rents and enters the regression with a very significant coefficient. Even controlling for this effect through income, however, diversity still plays a very important role in determining rents (coefficient equal to 0.90). While we take this as a potential sign that diversity has a positive amenity value (it actually shifts the free migration condition in Figure 3 to the left) we are concerned with the endogeneity of the income and population variables, and so we omit them in the rest of the analysis. To summarize, diversity has *positive and highly significant correlations* with both wages ($\beta_2 > 0$) and land rents ($\gamma_2 > 0$). These positive correlations can be interpreted as consistent with a dominant and positive effect of diversity on productivity.

Finally, as we have mentioned that employment and population are endogenous variables in the equilibrium of our model, let us consider another correlation that reinforces our interpretation of a dominant positive effect of diversity on productivity. The theoretical model makes clear (see equation (6)) that, in the presence of a positive productivity effect, the increase of diversity in a certain city shifts the local labor demand up, thus raising not only local wages but also local total employment. In contrast, a negative utility effect would be associated with higher wages but lower native employment. Table 4 reports the correlation between changes in diversity and changes in employment as well as the population of US cities between 1970 and 1990. If the labor supply curve had shifted up and the labor demand curve remained fixed, we should observe an increase in wages but a decrease in total employment caused by the outflow of US-born workers. The Table rather shows positive effects of diversity on both employment and population, consistent with the idea that there was no outflow of natives counterbalancing immigration. This is consistent with a dominant upward shift of labor demand as expected in the presence of a dominant positive productivity effect.

5.2. Checks of robustness

Our basic specifications for the wage (I and VI in Table 3) and rent (VII and IX in Table 3) regressions omit several variables that, in principle, could simultaneously affect local diversity, wages, and rents, thus creating spurious correlation. In so far as they change over time, the impacts of such omitted variables are not captured by city fixed effects. We have already discussed the potential roles of employment, income and population in the previous section. This section is devoted to testing whether the estimated

Table 4. Correlation between growth in diversity and in employment/population

Dependent variable:	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%.
Heteroskedasticity-robust standard errors are reported in parentheses.

effects of diversity are robust to the inclusion of other omitted exogenous variables. While our list of potential controls can not be considered exhaustive, we do include some important ones for which we can think of plausible stories that could generate spurious correlation. Table 5 reports the estimated effects of the diversity index (and its components) in the wage equation as we include additional controls. Table 6 presents analogous results for the rent regression. The coefficients in each row of Tables 5 and 6 arise from separate regressions. While it may be informative to discuss each regression in detail, we prefer simply to focus on the coefficients of interest; thus for the sake of brevity we comment only briefly on each specification. This section is meant to give the reader a general impression of the robustness of our estimates to a very ample range of controls, rather than to analyze in detail any one of the alternative specifications proposed.[16]

The positive effect of diversity on the wage of the US-born may simply be a result of the foreigners' measurable average education. Specifications (2) in Tables 5 and 6 include the average years of schooling of the foreign-born workers as an additional control variable in the wage and rent regressions respectively. While analyzing human capital externalities using average schooling has been a common practice (Rauch, 1993; Moretti, 2004), if workers with different schooling levels are imperfect substitutes, or if the distribution of their skills matters, then average schooling may not be a sufficient statistic to capture the presence of complementarity or externalities. The estimated effects of diversity on wages and rents remain significant and positive when we include this control. Interestingly, the effect of the average schooling of the foreign-born on the wages of the US-born (not reported) is not significant, while it is small and positive on the rents of the US-born. This result tells us that the simple average schooling of the foreign-born does not fully capture their true 'value.' Not only might the skill distribution of the foreign-born matter, but their abilities may be differentiated from (and complementary to) those of natives, even at the same schooling level. When we decompose the overall diversity (column 2 and 3 in the Tables) by including separately the share of foreign born and their diversity, we still find a significant and positive effect of the share of foreign born on both rents and wages, while the diversity of foreigners has a significant positive impact on wages but not on rents.

Another plausible (but spurious) reason for positive correlations between diversity and wages-rents may be that immigration responds to productivity and amenity shocks.

16 If the reader is interested in the details of each regression and in a more thorough discussion of each specification we suggest reading the working paper Ottaviano and Peri (2004b).

Table 5. 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	1.27**	0.57**	0.14*
	(0.30)	(0.11)	(0.08)
(2) Including schooling of foreign born	1.26**	0.56**	0.14*
	(0.38)	(0.16)	(0.09)
(3) Including share of out of state born	1.35**	0.58**	0.09
	(0.38)	(0.15)	(0.11)
(4) Including share of non whites	1.39**	0.66**	0.12
	(0.40)	(0.17)	(0.10)
(5) Including public spending on local services per capita	1.28**	0.63**	0.14*
	(0.38)	(0.17)	(0.09)
(6) Including public spending in education per capita	1.27**	0.65**	0.13
	(0.38)	(0.16)	(0.09)
(7) Including employment of white-US born males 40–50.	1.32**	0.67**	0.14
	(0.39)	(0.16)	(0.10)
(8) Including all of the above	1.43**	0.75**	0.10
	(0.40)	(0.18)	(0.08)
(9) Basic without CA, FL, NY	0.96**	0.23	0.21**
	(0.49)	(0.27)	(0.12)
(10) In changes 1990–1970 with state-fixed effects	0.85**	0.64**	0.02
	(0.31)	(0.17)	(0.12)
(11) Using wage of white-US born males 30–40 as dep. variable	1.20*	0.69*	0.04
	(0.37)	(0.14)	(0.10)

Dependent variable: ln average yearly wage to white, US born, males 40–50 years old expressed in 1990 US$. The coefficients in column 1 correspond to different regressions in each row. The coefficients in column 2 and 3 correspond to different regressions for each row.
(1) Basic: specification from Table 3 column I (for coefficient 1) and Column VI (for coefficients 2 and 3).
(2) Includes average years of schooling of foreign born.
(3) Includes the share of US-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 ln(Employment) of the group US-born, white males 40–50 years old.
(8) Includes all the variables in (1)–(7) together as controls.
(9) Excluding from the regression MSAs in the biggest immigrations states: CA, FL, NY.
(10) Regression in changes including 49-state fixed-effects.
(11) Uses the wage of the group white, US, born, males, 30–40 years old as dependent variable.
**Significant at 5%, *significant at 10%.
Heteroskedasticity-robust standard errors are reported in parentheses.s

In so far as we do not observe these shocks, we are potentially omitting the common underlying cause of changes in wages, rents and diversity. To address this issue we use two strategies. The first strategy, which we postpone implementing until Section 3, attempts to identify a variable correlated (or at least more correlated) with the share of foreign born and not otherwise correlated (or at least less correlated) with shocks to productivity or amenities. Then, it uses this variable as an instrument in the estimation. The second strategy, pursued here, exploits the fact that productivity shocks which attract workers into a city should attract the US-born and the

Table 6. 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	1.90**	1.13**	0.12
	(0.50)	(0.20)	(0.13)
(2) Including schooling of	2.00**	1.24**	0.14
foreign born	(0.59)	(0.23)	(0.15)
(3) Including share of out	1.98**	1.03*	0.22
of state born	(0.59)	(0.24)	(0.17)
(4) Including share of non	1.50**	0.96**	0.09
whites	(0.62)	(0.26)	(0.16)
(5) Including Public spending	1.93**	0.98**	0.22
on local services per capita	(0.59)	(0.25)	(0.16)
(6) Including public spending	1.92**	0.98**	0.22
in education per capita	(0.58)	(0.25)	(0.16)
(7) Including population of	1.50**	0.96**	0.08
white US-born males	(0.62)	(0.26)	(0.16)
(8) Including All of the above	1.69**	1.12**	0.07
	(0.60)	(0.27)	(0.16)
(9) Basic without CA, FL, NY	4.70*	1.23*	0.24*
	(1.20)	(0.27)	(0.16)
(10) 1990–1970	0.15	0.21	0.14
with state-fixed effects	(0.64)	(0.31)	(0.20)

Dependent variable: ln average monthly rent paid by white, US born, expressed in 1990 US$. The coefficients in column 1 correspond to different regressions in each row. The coefficients in column 2 and 3 correspond to different regressions for each row.
(1) Basic: specification from Table 4 column VII (for coefficient 1) and column IX (for coefficients 2 and 3).
(2) Includes average years of schooling of foreign born.
(3) Includes the share of US 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 the ln(population) of white US-born males.
(8) Includes all the variables in (1)–(7) together as controls.
(9) Excluding from the regression MSAs in the biggest immigrations states (CA, FL, NY).
(10) Regression in changes including 49 state fixed-effects.
**Significant at 5%, *significant at 10%.
Heteroskedasticity-robust standard errors are reported in parentheses.

foreign-born by the same degree. Therefore, the share of US-born citizens in each city coming from out of state (i.e. born in a different state) is a variable that should be correlated with the same local productivity and amenities shocks that attract foreigners.[17] Accordingly, its inclusion in the wage and rent regressions should

17 It may be the case, however, as argued by Borjas (2001), that the US-born move away from cities in which immigrants go because they look for different amenities or better wages. However, both our results shown in Table 4 (population increases where diversity increases) as well as recent studies by Card (2001) and Card and Di Nardo (2000) do not find evidence of this ‘displacement effect’.

significantly decrease the estimated coefficients β_2 and γ_2. Moreover, we should find a significant positive correlation between this share and the wage-rents of US-born citizens. Specification (3) in Tables 5 and 6 include the share of US-born citizens who were born out of state. Its coefficient (not reported) is not significant in either regression, while the effects of diversity and 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 experienced an unobserved positive shock, since that would have attracted both foreign and US-born workers. Interestingly, they also imply that their presence does not simply reveal that boom cities have attracted more talented people, since people of similar talent should respond similarly to the same shock.

Some sociologists have advanced the hypothesis that environments which are tolerant towards diversity are more productive and more pleasant to live in. Along similar lines, Richard Florida (2002a, 2002b) has argued that cities with larger numbers of artists and bohemian professionals are more innovative in high tech sectors. It is likely that part of our correlations may actually depend on this positive attitude of cities towards diversity. However, to show that there is something specific to the presence of foreign-born, we include in specification (4) of Tables 5 and 6 the share of US-born people identifying themselves as 'non-white.' Since we consider only US-born people, this index essentially captures the white-black composition of a city. The coefficients on this variable (not reported) turn out to be positive in the wage regression (0.20) and negative in the rent regression (−0.22). We may interpret these results as (weak) evidence of the aversion white US-born individuals feel living close to large non-white (US-born) communities. The standard errors however (in both cases around 0.2), render the estimated coefficients insignificant. As to the coefficients of the diversity index, they are still positive, significant (except in one case for the rent regression), and similar to previous estimates. Thus, in spite of the more ambiguous effect of ethnic diversity, diversity in terms of the country of birth maintains its importance.

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 whose quality of public services has improved in our period of observation may have experienced both an increase in the share of foreign born (possibly because they are 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 services per person in a city and on its breakdown across different categories, particularly in education. Specification (5) of Tables 5 and 6 includes overall spending by local government, whereas specification (6) includes spending on just education, a very important determinant of the quality of schools. The effect of public spending per person on rents (not reported) is positive in both specifications; however, its inclusion does not change the effects of diversity.

If different groups of workers are imperfect substitutes, then even among US natives the average wage of white males 40–50 years of age may be affected by their relative supply. While there is no clear reason to believe that the relative size of this group is correlated with the diversity of a city, it may be appropriate to control for the (log) employment of this group, and not just for total employment. The corresponding results are reported in specification (7) of Table 5, which shows that the coefficient of the diversity index is still equal to 1.32. Specification (7) of Table 6 considers instead the group of white US-born males as potentially competing for similar housing, and therefore includes the log of their population together with that of total population. This specification is very similar to specification (4), which includes the share of non-whites and produces similar estimates: 1.50 for the coefficient on diversity and 0.96 for the coefficient on the share of foreign born.

The most conservative check is specification (8), which includes together all the controls that are included separately in specifications (2) to (7). Reassuringly, the coefficient on the share of foreign-born is still positive, very stable, and significant in both regressions. The coefficient on the diversity index is also positive, very stable, and significant in the wage regression, while it turns out not significant in the rent regression.[18]

Specifications (9) and (10) of Tables 5 and 6 push our data as far as they can go. Specification (9) 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 a few highly diverse cities in those states generate the correlations of diversity with wages and rents. This turns out not to be the case. In the wage regression the coefficient on diversity decreases somewhat but remains both positive and significant. In the rent equation the coefficient on diversity grows larger but also becomes 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 low immigration states.

In Specification (10), rather than use city and year dummies, we use the differences of the basic variables between 1990 and 1970. We also include state fixed effects to control for differences in the 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. This is an extremely demanding specification as we are probably eliminating most of the variation needed to identify the results by 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.

We perform one more check in specification (11) of Table 5 in order to verify that our results survive when we consider groups that are more mobile across cities than 40 to 50 year-old workers. We estimate the wage equation using the average wage of white US-born males between 30 and 40 years of age. The coefficients on diversity and the share of foreign born are still significantly positive, equal to 1.20 and 0.69, respectively.

Finally, since our theoretical model shows that in equilibrium wages and rents are simultaneously determined (see equations (11) and (12)), thus implying correlation between the unobservable idiosyncratic shocks to wages $\varepsilon_{c,t}$ and rents $e_{c,t}$, we can increase the efficiency of our estimates by explicitly accounting for such correlation, and estimate a seemingly unrelated regression (SUR). While OLS estimates are still consistent and unbiased even when $\varepsilon_{c,t}$ and $e_{c,t}$ are correlated, SUR estimates are more efficient. The estimated coefficients are virtually identical to those estimated in Table 5 and 6. For sake of brevity we do not report the results here.[19]

18 Some authors (see e.g. Sivitanidou and Wheaton, 1992) have argued that 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. This effect, however, should be captured by our local public goods measures.

19 The results of SUR estimations are available in Ottaviano and Peri (2004b).

In summary, most wage and rent regressions yield positive and significant coefficients for both the diversity index and the share of foreign born. The diversity of the foreign born also has a positive effect but this effect is less often significant. We do not find any specification such that the coefficients on the diversity variable are simultaneously not significant in both the wage and the rent regressions. Moreover, each single estimate delivers positive estimates of diversity on wages and rents of natives. Therefore, our identification (13) allows us to conclude that *no specification contradicts the hypothesis of a positive productivity effect of diversity*.

5.3. Endogeneity and instrumental variables

Short of a randomized experiment in which diversity across cities is changed randomly, we cannot rest assured that our correlations reveal any causal link from diversity to wages and rents. Nonetheless, some steps towards tackling the endogeneity problem can be taken using instrumental variables (IV) estimation. Our instruments should be correlated with the change in the diversity of cities between 1970 and 1990, and not otherwise correlated with changes in wages and rents. We construct our main instrument building on the fact that foreigners tend to settle in 'enclaves' where other people from their country of origin already live (Winters et al., 2001; Munshi, 2003). Following Card (2002) and Saiz (2003b) we construct the 'predicted' change in the number of immigrants from each country in each city during the observed period. The predicted change is based on the actual shares of people from each country in each city at the beginning of the period, and the total immigration rate from each country of origin to the US during the whole period. By construction the 'predicted' change does not depend on any city-specific shock during the observed period. We then observe that the stocks and flows of immigrants tend to be larger in cities that are closer to important 'gateways' into the US. By contrast, the stocks of the native born and their changes over time are much less dependent on their proximity to these gateways. Therefore, as additional instruments, we also add the distance of a city from the main gateways into the US after having tested for the exogeneity of these instruments. The inclusion of more instruments, as long as they are exogenous, should improve our estimates while still correcting for the potential endogeneity bias. We now describe the instruments and the estimation results in the following two sections.

5.3.1. Shift-Share methodology

We construct our main instrument by adopting the 'shift-share methodology,' used by Card (2001) and more recently by Saiz (2003b), to migration in MSA's. Immigrants tend to settle, at least initially, where other immigrants from the same country already reside (immigration enclaves). Therefore, we can use the share of residents of an MSA in 1970 for each country of birth, and attribute to each group the growth rate of that group within the whole US population in the 1970–1990 time period. In so doing we compute the predicted composition of the city based on its 1970 composition and attribute 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 a 'predicted' diversity index for each city in 1990.

Let us use the notation introduced in Section 3.1, where $\left(CoB_i^c\right)_t$ denotes the share of people born in country i among the residents of city c in year t. Hence,

$(CoB_i)_t = \sum_c (CoB_i^c)_t$ is the share of people born in country i among US residents in year t. Between 1970 and 1990 its growth rate is:

$$(g_i)_{1970\text{–}90} = \left[(CoB_i)_{1990} - (CoB_i)_{1970}\right] / (CoB_i)_{1970} \tag{16}$$

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

$$(\widehat{CoB_i^c})_{1990} = \left(CoB_i^c\right)_{1970}\left[1 + (g_i)_{1970\text{–}90}\right] \tag{17}$$

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

$$\widehat{div}_{c,1990} = 1 - \sum_i (\widehat{CoB_i^c})^2_{1990} \tag{18}$$

As the attributed diversity for each city in 1990 is built using the city's share in 1970 and the 1970–1990 national growth rates of each group, this value is independent from any city-specific shock during the period.

Tables 7 and 8 present the results of the IV estimation of the wage and rent regressions. Relative to previous regressions, some adjustments in the grouping of countries of birth are needed. This is because as we input the shares in 1990 based on the initial shares in 1970, we need to identify the same countries of origin across census years. This is achieved by allocating more than one country of birth to the same group, as some countries have disappeared or changed during the period. In so doing, we follow the classification adopted by Card (2001) and described in the data appendix.

In Tables 7 and 8, column 1 reports the OLS estimates of the basic specification in which we control for schooling using the change in average years of schooling in the city (Δ schooling). The point estimates of the OLS specification are very similar to our previous estimates (Table 3, columns I and VII), confirming that the reclassification

Table 7. Wage regression. IV estimation, instrument: shift-share imputed diversity

Dependent variable : Δln(wage)	1 OLS in differences	2 Controlling for initial average wage	3 IV	4 IV without CA-FL-NY
ΔSchooling	0.11**	0.11**	0.11**	0.10**
	(0.01)	(0.01)	(0.01)	(0.01)
Δ(diversity)	1.27**	1.43**	0.98**	0.99*
	(0.38)	(0.39)	(0.50)	(0.60)
R^2	0.34	0.36	0.35	0.33
Observations	160	160	160	145
		First stage regression		
Shift-share constructed diversity	n.a.	n.a.	0.51**	0.21**
			(0.05)	(0.04)
Partial R^2	n.a.	n.a.	0.31	0.17

Dependent variable: change between 1970 and 1990 in ln average yearly wage of white, US born, males, 40–50 years, expressed in 1990 US $.
Instrumental variable: imputed change in diversity index and share of foreign born, using the shift-share method described in the text.
**Significant at 5%, *significant at 10%.
Heteroskedasticity-robust standard errors are reported in parentheses.

Table 8. Rent regression. IV estimation, instrument: shift-share imputed diversity

Dependent variable : Δln(rent)	1 OLS in differences	2 Controlling for initial average rent	3 IV	4 IV, Without CA-FL-NY
Δ(diversity)	1.97**	2.07**	2.60**	3.29**
	(0.60)	(0.65)	(0.96)	(1.50)
R^2	0.07	0.12	0.10	0.12
Observations	160	160	160	145
		First stage regression		
Shift-share constructed diversity	n.a.	n.a.	0.51**	0.21**
			(0.05)	(0.04)
Partial R^2	n.a.	n.a.	0.23	0.11

Dependent variable: Change between 1970 and 1990 in logged average yearly rent of white, US-born, aged 16–65, expressed in 1990 US$.
Instrumental variable: imputed change in diversity index and share of foreign born, using the shift-share method, described in the text.
**Significant at 5%, *significant at 10%.
Heteroskedasticity-robust standard errors are reported in parentheses.

by country groups has only small effects. In column 2, as we are running the specifications in differences (rather than in levels with fixed effects), we also check that the implicit treatment of long-run effects as equilibrium effects is appropriate. In particular we include the initial values of average wages and rents (coefficients on those variables are not reported), in order to control for the possibility that cities were not at a long-run equilibrium at the beginning of the period (1970), so that their dynamic behavior exhibits 'conditional convergence'. The estimated effects of diversity do not change much, and are statistically not different from the previous estimates.

As for the IV estimates of columns 3 and 4, we notice that the first stage regressions (of the endogenous measure of diversity on the instrument) imply that the imputed diversity indices are good predictors of the actual ones, explaining 31% of their variation (orthogonal to the other regressors) when all states are included. The exclusion of large immigration states, however, reduces significantly the partial R^2 of the first stage regression to 17%.

The estimated effect of diversity on wages is reported in column 3 of Table 7. Its value (0.98) is close to the OLS estimate and significantly positive. When we exclude the high-immigration states (column 4 of Table 7), the effect of diversity is estimated to be positive but significant only at the 10% confidence level. However, the main problem encountered when we exclude California, Florida and New York is that, as just mentioned, the instruments lose much of their explanatory power (the partial R^2 of the excluded instruments drops to 0.17). Therefore, insignificance is mostly driven by large standard errors, rather than by evidence of any endogeneity bias (i.e., changes in point estimates).

Columns 3 and 4 in Table 8 show that the rent regression exhibits a similar qualitative pattern but sharper results. Using the shift-share instruments, the diversity index has a positive and significant effect in each specification. Including all states, the IV estimates are 30% higher than the OLS estimates (although, due to the large standard error we cannot reject the hypothesis that they are equal). When we exclude California, Florida, and New York (specification 4 of Table 8), both the estimate and the standard error increase significantly. The point estimates of the effect of diversity are still firmly in the positive range. Somewhat surprising (possibly driven by the exclusion of some 'perverse' outliers such as Miami, see Figure 2) is the very large (and imprecisely estimated) effect of diversity on rents in this specification.

5.3.2. Gateways into the US

We can increase the set of instruments by noting the fact that immigrants tend to enter the US through a few 'gateways,' or through the border. As a consequence, the total number of foreign born in city c at time t, F_{ct}, as well as the total increase in foreign born in city c, ΔF_{ct}, depend negatively on the distance from the closest gateway. As long as the total number of US-born residents in a city, N_{ct}, does not depend (or depends to a lesser extent) on that distance, we have that both the share of foreign born, $F_{ct}/(F_{ct}+N_{ct})$, and its change are negatively correlated with the distance from the immigration gateways into the US.

Each year the US Office of Tourism publishes the percentage of inbound travellers by point of entry. Looking at this data for the eighties, we see that the three main gateways were New York, Miami, and Los Angeles. About 30% of foreign (immigrant and non-immigrant) travellers entered the US through the airports and ports of these cities. Moreover, due to the benefits of networks, the costs of travelling, and the costs of spreading information, immigrants were more likely to settle in cities closer to these gateways. 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 the natural place of entry into the US. Thus, as before, cities closer to these borders were more likely to receive Canadian or Mexican immigrants during the 1970–1990 period.

These considerations suggest the use of the overall distance of a city from the main gateways into the US (New York, Miami, Los Angeles, and the US borders with Canada and Mexico) to instrument for its diversity index (heavily dependent on the share of foreign-born). This distance should be negatively correlated with diversity but not with shocks to wages and rents.

This strategy is certainly open to criticism. If the three main gateways (New York, Miami, and Los Angeles) or the region of the US-Mexican border experienced above average growth during the time period considered, then positive spillover effects on nearby cities could attract foreigners. As a result, the distance of a city from these gateways would be negatively correlated with the increases in wages and rents because of a 'boom city' effect rather than a positive effect from diversity. This criticism, however, does not apply to the 'predicted diversity' constructed in the previous section. As we are confident of the 'exogeneity' of one instrument (the 'predicted diversity'), when using additional instruments (distance from gateways) we can test for their exogeneity[20]. We find that the variables that do not fail the exogeneity test jointly are 'predicted diversity', distance from NY, distance from LA and distance from Miami. We had to drop the distance from the border variable, as it failed this exogeneity test.

20 The exact form of our test of exogeneity can be find in Woolridge (2001), 124–125. Intuitively the test checks wether the restriction that excludes the extra-instruments from the second-stage regression is rejected or not by the data. If it is not rejected the assumption of exogeneity stands.

Table 9. Wage regression. IV estimation, instruments are distance from 'Gateways' and imputed diversity

Dependent variable : Δln(wage)	1 IV	2 IV with state effects	3 IV, without CA-FL-NY
ΔSchooling	0.11**	0.11**	0.11**
	(0.01)	(0.02)	(0.01)
Δ(Diversity)	1.50**	0.68**	1.91**
	(0.39)	(0.33)	(0.54)
State fixed effects	No	Yes	No
R^2	0.35	0.63	0.30
Observations	160	160	144
	First stage regression		
Shift-share constructed diversity	0.44**	0.44**	0.30**
	(0.04)	(0.04)	(0.04)
Ln(distance from LA)	−0.01**	−0.01**	−0.01**
	(0.001)	(0.001)	(0.002)
Ln(distance from NY)	−0.005**	−0.005**	−0.006**
	(0.0008)	(0.0008)	(0.0007)
Ln(distance from Miami)	−0.01**	−0.01**	−0.004**
	(0.001)	(0.001)	(0.002)
Partial R^2	0.71	0.51	0.46

Dependent variable: change between 1970 and 1990 in ln average yearly wage of white, US-born, males, 40–50 years, expressed in 1990 US$.
**Significant at 5%, *significant at 10%
Heteroskedasticity-robust standard errors are reported in parentheses.
Test of over-identifying restrictions, from Woolridge (2001) pp. 124–125, cannot reject the joint exogeneity of instruments at the 5% confidence level. The value of the test statistic is 3.2 for the first specification, 4.5 for the second and 3.7 for the third. The statistic is distributed as a chi-square with 3 degrees of freedom under the null hypothesis of no Instrument included in the second stage equation.

Tables 9 and 10 report the first and second stage estimates of the described IV regressions using wages and rents, respectively, as the dependent variable. Column 1 of Table 9 shows the basic specification of the wage regression; column 2 includes 48 state fixed-effects; column 3 excludes the biggest immigration states. Similarly column 1 of Table 10 includes the basic specification while column 2 and 3 exclude from the regression coastal cities and cities in California, Florida and New York as a check for potential outliers driving the results. The first stage regressions confirm that our excluded instruments are excellent: in the first stage they explain about 70% of the variation in diversity that is orthogonal to the other regressors. Even including state effects, more than 50% of the residual variation in diversity is still explained by the instruments. This increases the power of instrument, relative to Table 7 and 8 and may result in more precise estimates.

The estimates of specification 1 (Table 9 and 10) confirm that the effects of diversity on wages and rents are positive and large. The estimated coefficient is significant and very large for wages (1.50) as well as for rents (1.48). Moreover, the IV estimates of the effect on wages are somewhat higher than the OLS ones; hence we are reassured that no significant (endogeneity-driven) downward OLS bias exists. For the wage regressions we obtain a positive and significant effect of diversity when controlling for 48 state fixed effects (specification 2 of Table 9) and when eliminating coastal cities (specifications 3 of Table 9). The last specification has quite large standard errors, but it certainly

Table 10. Rent regression. IV estimation, instruments are distance from 'Gateways' and imputed diversity.

Dependent variable : Δln (rent)	1 IV	2 IV non-coastal cities	3 IV, without CA-FL-NY
Δ(Diversity)	1.48**	5.50**	4.70**
	(0.61)	(2.31)	(1.04)
State fixed effects	No	No	No
R^2	0.13	0.10	0.12
Observations	160	160	144
	First stage regression		
Shift-share constructed diversity	0.44**	0.23**	0.30**
	(0.04)	(0.05)	(0.04)
Ln(distance from LA)	−0.01**	−0.005**	−0.01**
	(0.001)	(0.001)	(0.002)
Ln(distance from NY)	−0.005**	−0.004**	−0.006**
	(0.0008)	(0.0008)	(0.0007)
Ln(distance from Miami)	−0.01**	−0.01**	−0.004**
	(0.001)	(0.001)	(0.002)
Partial R^2	0.71	0.38	0.46

Dependent variable: change between 1970 and 1990 in ln average monthly rent paid by white, US-born, expressed in 1990 US$.
**Significant at 5%, *significant at 10%.
Heteroskedasticity-robust standard errors are reported in parentheses.
Test of over-identifying restrictions, from Woolridge (2001) pp. 124–125, cannot reject the joint exogeneity of instruments at the 5% confidence level. The value of the test statistic is 4.8 for the first specification, 7.2 for the second and 4.5 for the third. The statistic is distributed as a chi-square with 3 degrees of freedom under the null hypothesis of no instrument included in the second stage equation.

reinforces our thesis that the foreign-born have a positive effect in non-coastal cities as well. As for the rent regressions, the share of foreigners once again has a positive and significant effect in specifications 2 and 3 of Table 10 (excluding coastal cities and excluding the largest immigration states). Again, somewhat oddly, and probably due to the elimination of some outliers, the estimated effect on rents increases significantly in specifications 2 and 3.

All in all the results using shift-share instruments seem to confirm very strongly the positive effect of diversity on wages and rents of natives. In particular, considering all the IV regressions, we find no specification in which the coefficients of diversity are not significant in either the wage or rent equations. Moreover the point estimates are always robustly positive (although sometimes they are not very precise due to instrument weakness). Thus, on the basis of the discussion in subsection 2, we can conclude that our data support the hypothesis of a positive productivity effect of diversity with *causation running from diversity to productivity of US workers*.

6. Discussion and conclusions

We have looked at US metropolitan areas as a system of open cities in which cultural diversity can affect the productivity and utility of natives. In principle, the effects of diversity can be positive or negative. We have considered a simple model that handles all possible cases (i.e. positive or negative effects on productivity and utility), and we have designed a simple identification procedure to figure out which case receives empirical support based on cross-city wage and rent variations. We have showed that higher wages and higher rents for US natives are significantly correlated with higher diversity. This result has survived several robustness checks against possible alternative explanations based on omitted variables and instrumental variables estimation.

Given our identification procedure, these findings are consistent with a dominant positive effect of diversity on productivity: *a more multicultural urban environment makes US-born citizens more productive*. To the best of our knowledge, in terms of both data and identification procedure, our results are new. We need to add two caveats, however, to these conclusions. First, while we are confident that the identified positive correlation between diversity and wage-rents is a robust feature of the data, our interpretation of a positive effect of diversity on productivity is not the only possible one. A plausible, and not less interesting one, is that spatial selection of US born residents in cities with high or low diversity may reflect some of their characteristics. For instance, people with higher education, higher international experience, and higher exposure to culture and news may be more appreciative of diversity. They may also be different from other US natives in several characteristics that are related to productivity. If this is true, 'tolerant' cities are more productive due to the characteristics of US-born residents rather than to the 'diversity' of these cities. Our current and future research is proceeding in the direction of analyzing this selection effect better and trying to determine which factor (diversity or tolerance) is more relevant for productivity (in fact both effects are likely to play important roles).

Secondly, even assuming the existence of a positive effect of foreign-born residents on the productivity of US natives, we have not yet opened the 'black box' to analyze theoretically and empirically what the channels are through which that effect works. The complementarity of skills between the US and foreign born seems a very promising avenue of research. 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. Another promising avenue is that foreign-born workers may provide services that are not perfectly substitutable with those of natives. An Italian stylist, a Mexican cook and a Russian dancer simply provide different services that their US-born counterparts cannot. Because of a taste for variety, this may increase the value of total production. We need to analyze more closely the effects in different sectors and on different skill groups in order to gain a better understanding of these channels. Overall our findings look 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 year) 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 considered only the countries of origin in which was born at least 0.5 % of the foreign-born working age population. We obtained 35 groups for 1970 as well as for 1990.

We used the Variable 'Salary and Wage' to measure the yearly wage income of each person. We transformed the wage in real 1990 US$ by deflating it with the national GDP deflator. The years of schooling for individuals 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's (1994) correspondence Table 4. Average rents are calculated using 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.

A.2 Grouping by country of birth

In Tables from 1 to 8 we consider the diversity index constructed using 35 countries of origin of immigrants which top the list of all countries of origin plus a residual group called 'others'. These account for more than 90 % of all foreign-born, both in 1970 and 1990, and a country that is not in this list supplies at most 0.5 % 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, USSR, 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 9 and 10, in order to have the same groups 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). This is the list: Mexico, Caribbean Countries, Central America, China-Hong-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.

Table A1. 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

Table A1. *Continued*

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