Market potential and city growth: Spain 1860–1960

Rafael González-Val (Universidad de Zaragoza-IEB)
Daniel A. Tirado-Fabregat (Universitat de València)
Elisabet Viladecans-Marsal (Universitat de Barcelona-IEB)

Abstract:

In this paper, we employ parametric and nonparametric techniques to analyse the effect of market potential on the structure and growth of Spanish cities during the period 1860–1960. Even though a few attempts have been made to analyse whether market potential might influence urban structures, this period is especially interesting because it is characterised by advances in the economic integration of the national market together with an intense process of industrialisation. By using an elaborated measure of market potential at the city level, our results show a positive influence of this market potential on city growth, although this influence is heterogeneous over time. Only changes in the market potential from 1900 have a significant effect on population growth.

Keywords: market potential, urban structure, city growth, economic history

JEL codes: R0, N9, O18, N64, F14

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1. Introduction

A growing body of empirical studies analyse the incidence of market potential on the geographical distribution of population and economic activity, migratory patterns, wage levels and differences in regional income.[1] However, to date, few studies have examined the relation between market potential and the patterns of city size growth. Recently, a number of papers have introduced market size effects in their explanations of the geographical distribution of cities and of their relative sizes. Indeed, in a recent survey, Redding (2010) points out that it might be interesting to reconcile new economic geography (NEG) models with the findings of the urban economics literature regarding the distribution of population size and population growth patterns. As Krugman (1991) claims, two types of factors can be considered to be determinants of city growth: *first nature* factors, which are related principally to geography (climate, costal location, access to natural resources, etc.) and which influence city growth in their early stages; and *second nature* factors, which are related to agglomeration economies and increasing returns of scale.

Here, a part of the literature has considered change in market potential to be a good proxy for agglomeration economies. However, the direction of the influence of market potential on city growth is unclear. The trade theory literature (Davis and Weinstein, 2002; Hanson, 2005) concludes that greater market potential should foster growth, the rationale being that nearby cities offer a larger market and, hence, more possibilities of selling products. By contrast, location theory (Fujita et al., 1999) and hierarchy models (Dobkins and Ioannides, 2001) suggest that increasing market potential could affect city growth negatively, the rationale being that the forces of spatial competition separate the larger cities from each other, so the bigger a city grows, the smaller its neighbouring cities will be. Finally, it is interesting to note that the effects of market potential on city size may differ depending on the initial size of the city.

Although there is a sizeable body of theoretical research developing models around these factors, the empirical evidence remains limited. In recent years, various papers have specifically analysed the incidence of market potential on city growth. For US cities, Black and Henderson (2003) analyse the determinants of population growth from a long-term perspective. Henderson and Wang (2007) analyse the influence of market potential on

[1] See Redding (2010) for a survey on this literature.

population growth for the US metropolitan areas between 1960 and 2000. Au and Henderson (2006) adopt a similar approach in their analysis of Chinese cities during the nineties. Finally, da Mata et al. (2007) analyse the determinants of city growth for Brazilian cities. The results of these papers seem to confirm the incidence of the increase in market potential on city growth.

This paper is conducted in line with the preceding studies focusing the analysis on the effects of the economic integration and industrialisation of the Spanish economy on the evolution of the urban system during the period 1860–1960. The hypothesis we test is the following: the geographical distribution and relative size of Spanish cities were historically determined by the location fundamentals of each territory. However, when the Spanish market began to be integrated, there was an increase in the concentration of the population in a small number of cities. This concentration could explain the increase in the inequalities in relative city sizes. In other words, in a context in which manufacturing activities were acquiring greater weight in the economy, the construction of the transport network and integration of the markets would have favoured, from the middle of the 19th century, the agglomeration of economic activities and this could have been the basis for the changes in long-run city patterns. Basically, the urban systems that prevailed before and after market integration and industrialisation were quite distinct. Our results lend support to these hypotheses and confirm that market potential had a clear influence on these processes, although this influence is heterogeneous over time. More specifically, a 1% increase in market potential implied an average 0.10% increase in the population of the city, although this effect is nonlinear depending on the initial market potential level and changes over time. Thus, only changes in the market potential from 1900 have a significant effect on population growth.

The contribution of this paper is twofold. On the one hand, we exploit the long-term historical episode of growth and economic integration that took place in Spain from the middle of the 19th century until the 1960s. Thus, we are able to study the determinants of city growth throughout the whole process of the market integration of the Spanish economy and at a time when the Spanish urban system was undergoing an intense transformation characterised by the concentration of the population in a small number of cities and with a clear impact on their relative sizes. Nevertheless, the study of the determinants of the relative growth of Spanish cities along the years 1860–1960 provides a sound contribution to the empirical NEG literature. As pointed out by Redding (2010), the study of the effects of market access on the regional distribution of economic activity or population faces an important empirical problem. It is difficult to disentangle the effects of market access from other determinants of regional or urban growth such as locational fundamentals, meaning that the results of a great bulk of empirical analyses are subject to a problem of the indeterminacy of the causality of the found relationships. The literature has suggested the analysis of these types of relations in the context of exogenous changes in the relative market size of territories. Examples of this type of approach can be found in Hanson (1996), Wolf (2007), Davis and Weinstein (2002) and Redding and Sturm (2008). In this sense, therefore, the case study of the Spanish experience during this period allows analysing whether the construction of new transport infrastructure, as well as changes in trade policy, that exogenously affected the market potential of Spanish territories (Martínez-Galarraga, 2012), ended up shaping relative city growth.

On the other hand, following Black and Henderson (2003), Ioannides and Overman (2004) and Bosker et al. (2008), we test the importance of market potential on city growth. Nevertheless, in our empirical approach, we do not use the distance-weighted sum of the population of all other existing cities as a proxy of city market potential. We depart from the work by Martínez-Galarraga (2010) that presents an empirical measure of regional market potential for the inland Spanish NUTS3. This measure considers a new set of historical GDP estimates for Spanish regions and historical transport costs as well as the changes they underwent during the process of the economic integration of the Spanish economy. From this starting point, we compute a measure of city market potential for each of the 266 to 1030 Spanish cities in our database. Therefore, we distribute each NUTS3 market potential across all the cities belonging to each region according to the city population share in the whole regional population and the relative size of the surrounding cities within the region.[2]

The remainder of the paper is organised as follows. In section 2, we describe the economic integration process that took place in Spain from 1860 to 1960 and review the main evidence from the literature regarding its economic effects. In section 3, we analyse the evolution of Spanish city size distribution from 1860 to 1960. In section 4, we present our

[2] We use a measure of real transport costs. Combes and Lafourcade (2005) conclude that this is a better approach.

data. In section 5, we describe the empirical analysis and present and discuss the results obtained. Section 6 concludes.

2. Market integration and economic agglomeration in Spain, 1860–1960

From a long-term perspective, Spain's internal market integration received a major push in the middle of the 19th century. Prior to this date, Spanish regions had relatively independent economies. Barriers to interregional trade and the movement of capital and labour were ubiquitous: local tariffs and regulations on domestic commerce were widespread; weights and measures differed across regions; transport costs were very high due to low public investment in transport infrastructures and the particular geography of Spain, which lacked an extensive water transport system; economic information moved slowly across regions; the banking system was underdeveloped; and many regions had their own currencies (although they were all based on a bimetallic monetary system). As a consequence, regional commodity markets were scarcely integrated and the prices of production factors differed markedly from one region to another.[3]

The 19th century political reforms strengthened property rights and reduced transaction costs that interfered in economic relations and impeded the free movement of goods and factors within Spain. Importantly, reforms eliminated the main restrictions on trade (including tariffs and domestic customs), suppressed guilds and the *Mesta* (a medieval association of cattle farmers), disentailed land (*desamortización*), abolished entailed states (*mayorazgos*) and unified the system of weights and measures that had hitherto varied from region to region. In addition, a decree in 1868 by treasury minister Laureano Figuerola unified Spain's monetary system, which was henceforth founded on a single currency: the peseta (in 1864, 84 different coins remained in circulation). Besides monetary unification, the banking system advanced in several ways. It began to modernise during the early 1840s, completing the process in 1874 when the Echegaray Decree abolished the plural system based on various banks of issue and granted monopoly to the Bank of Spain. The Bank of Spain had also begun to open branches in provincial capitals.

Improvements in transport systems, particularly the completion of Spain's railway network, prompted the creation of a national market for most major commodities during the second

[3] See, for example, Ringrose (1996) for further details.

half of the 19th century. Although the construction of the Spanish railroad network started in 1848, it was not until 1861 when the main inland territories in Castile were connected to the coastal provinces in the North and on the Mediterranean coast. According to calculations reported by Herranz (2006), the introduction of the railway in 1878 meant a massive 86% reduction in transport costs. In addition to these two factors, successive political reforms of the 19th century upheld property rights, eliminated tariffs and local restrictions on home commerce and safeguarded the mobility of people and capital. These measures were implemented in three main waves: during the Liberal Revolution (1836–1840), the two-year Progressive period (1854–1856) and the six-year Democratic period (1868–1874).

The outcome was the gradual integration of the national market for goods for the main traded products, an integration characterised by the convergence in regional prices. Various studies have proven the gradual convergence of regional grain prices from the beginning of the 18th century until its culmination in the second half of the 19th century (Peña and Sánchez-Albornoz, 1983). In addition, the integration of the markets for capital and labour underwent marked advances as well. In the case of capital markets, the main events that affected the monetary and banking systems favoured a reduction in interest rates differentials across regions. In particular, Castañeda and Tafunell (1993) show that interregional short-term interest rate differentials registered an intense decline after 1850. Lastly, Spain's labour market integration, measured in terms of disparities in regional real wages across regions, has also been extensively analysed. In this respect, Rosés and Sánchez-Alonso (2004) show that PPP-adjusted rural and urban wages converged across different locations prior to World War I despite low rates of internal migration.[4]

The Spanish Civil War and first years of Franco's regime acted as a brake on Spanish growth and its national economic integration. The regulation of markets for goods and factors of production combined with government control of the prices and quantities of final and intermediate goods, energy, capital markets and wages reduced the mobility of factors and resources. The movement of capital across regions slowed and labour migration came to a halt after an initial period of growth in the 1920s (Silvestre, 2005). Likewise, the absence of investment in infrastructure did little to reduce transport costs during the 1940s and early 1950s. The economic liberalisation and stabilisation measures

[4] A more detailed description of these processes can be found in Rosés et al. (2010).

introduced during the decade of the fifties, however, favoured the transition of the Spanish economy toward a new phase of economic development that would last until the oil crisis.

Recent studies have attempted to analyse the effects of this process of long-term economic integration and growth on the distribution of industry across Spanish regions.[5] Rosés (2003) and Tirado et al. (2002) provide new empirical evidence confirming that, from the second half of the 19th century until the outbreak of the Spanish Civil War, there was a marked increase in the geographical concentration of industry. In addition, both studies stress that this long-term evolution was in line with predictions emanating from NEG models. This strand of the literature suggests that the reduction in transport costs in the presence of scale economies in industrial activities results in the geographical concentration of industry and that production agglomerates in locations that enjoy the best access to markets. In other words, new evidence regarding the evolution of the geographical concentration of industry in Spain in the period that extends from 1860 to 1960 points to the fact that the relative market access of Spanish regions could act as an important explanatory factor of industrial agglomeration geography.

Spanish economic growth and market integration also favoured the increasing concentration of population across regions. In fact, the Gini index for regional population (at a NUTS3 scale) grew steadily from 0.266 in 1860 to a value of 0.402 in 1960.[6] Several studies have also explored the economic factors underpinning this process. First, in line with the empirical proposals made in the NEG literature, Paluzie et al. (2009) follow Crozet (2004) to demonstrate the existence of a direct relationship between the location decisions of migrants and market potential of host regions during the two main waves of internal migration in Spain in the 1920s and 1960s.[7] Second, Ayuda et al. (2010) analyse the patterns of the geographical distribution of the population in Spain from the 18th century onwards. They report that in the pre-industrial period, when agriculture was the predominant activity, *first nature* advantages determined the distribution of the population across Spanish provinces as climatic and topographic conditions had a direct impact on agrarian productivity. As a result, these natural conditions provided some locations with an

[5] Most of the empirical contributions to the Spanish case discussed below adopt the standard empirical methodologies developed for the analysis of NEG models. Redding (2010) offers a recent survey of these empirical methods.
[6] Ayuda et al. (2010).
[7] Kancs (2005) also makes use of this approach to analyse the determinants of migratory flows in the European Union.

initial advantage. However, the authors conclude that market integration in a context of industrialisation strengthened this pattern. From 1900 onwards, *second nature* geography, linked to increasing returns and relative access to regional markets, reinforced the process of the spatial agglomeration of the population.[8]

Summing up, this empirical literature records that industrial production and the population in Spain agglomerated parallel to the long-term process of development and market integration. Moreover, in line with the hypotheses derived from the NEG literature, differences in regional market access acted as a key factor in explaining the geography of this increasingly agglomerated economy. In line with these conclusions, the sections that follow are devoted, first, to presenting new evidence regarding the changes experienced in the Spanish urban system during this long-term process of economic development and market integration; and, second, to analysing the role played by differences in the market access of Spanish cities as a factor that accounts for their relative growth.

3. Changes in the Spanish urban system: The evolution of the city size distribution

This section analyses the evolution of Spain's city size distribution from 1860 to 1960. Other studies have examined this distribution, above all during the 20th century, identifying a divergent pattern of growth in city sizes during the period 1900–1970 (see Lanaspa et al., 2003, for a good example of this). Here, we seek to add to this literature by offering new empirical evidence dating back to 1860. We estimate Pareto exponents and empirical density functions. Our results show that from 1860 to the beginning of the 20th century, the city size distribution remained stable, but after that date, a process of divergent growth that increased inequality within the distribution is identified.

Our geographical unit of reference is the municipality (local government areas), the smallest spatial subdivisions in Spain's administrative system, which cover the whole territory and include all the country's population. Our population data are drawn from the 1860 census and thereafter from the decennial censuses conducted since 1900. Reher (1994) provides population data for 1860, while for all the other decades, we use data from the Spanish official statistics institute, the censuses of the *Instituto Nacional de Estadística* (INE - www.ine.es).

[8] In a similar vein, Goerlich and Mas (2009) also study the long-term determinants of the agglomeration of the population in Spain.

Herein Table 1

Table 1 shows the number of cities by period and their corresponding descriptive statistics. We impose a minimum population threshold of 5,000 inhabitants in each period since the smallest cities can hardly be considered to be urban (one of the particular features of the Spanish city system is the high number of small rural towns). Furthermore, until the middle of the 20th century, a considerable part of the country's employment was concentrated in the agriculture sector (38.7% in 1960 according to OECD data), so metropolitan structures only really began to emerge in the second half of the century. Figure 1, which plots two maps showing the spatial distribution of the municipalities in our samples in 1860 and 1930, shows that there was a sizable entry of new cities in the distribution by this later date. In 1860, most of the cities were located in Andalusia, the southernmost region of Spain, but several decades after, new cities had emerged in the centre and in the northwest of Spain.

Herein Figure 1

A standard way to analyse the evolution of the city size distribution involves fitting a Pareto distribution to the data (Cheshire, 1999; Gabaix and Ioannides, 2004). Let S_i be the size (population) of city i and R_i its corresponding rank (1 for the largest, 2 for the second largest and so on). We define the relative size of the *i*th city, s_i, as the quotient between the city's population and the contemporary average,

$$s_i = \frac{S_i}{\bar{S}} = \frac{S_i}{\frac{1}{N}\sum_{i=1}^{N} S_i}, \tag{1}$$

where N is the sample size. A power law (Pareto distribution) links city size and rank as follows: $R_i(s_i) = As_i^{-a}$, where A is a constant and a is the Pareto exponent. Zipf's law is an empirical regularity, appearing when the Pareto exponent of the distribution is equal to unity ($\hat{a} = 1$) and which means, when ordered from largest to smallest, the size of the second city is half that of the first, the size of the third is a third of the first and so on. Moreover, the greater the coefficient, the more homogeneous the city sizes.

Herein Figure 2

By taking logs, we obtain the logarithmic version usually estimated by OLS. We apply the specification proposed by Gabaix and Ibragimov (2011), subtracting $1/2$ from the rank to obtain an unbiased estimation of a:

$$\ln\left(R_i - \frac{1}{2}\right) = b - a \ln s_i + \varepsilon_i, \qquad (2)$$

where ε_i is the error term. We estimate Equation (1) by OLS for our sample of cities in the different periods from 1860 to 1960. Graph (a) in Figure 2 shows the results,[9] which demonstrate that the distribution remained stable from 1860 to 1900. Further, the estimated coefficients are greater than one, indicating that city sizes were homogeneous. Since the beginning of the 20th century, the estimated values of the Pareto exponent tend to fall, indicating a process of divergent growth in Spanish cities (Lanaspa et al., 2003). However, the exponent is always greater than one, rejecting Zipf's law. Graph (b) shows the results considering a balanced panel of the 262 municipalities existing in 1860 with a population above the minimum threshold, not allowing for the entry or exit of cities in the sample. The pattern for these cities is similar to that of the whole sample: the Pareto exponent is stable until 1900, when it begins to decrease. The only difference is that for this sample of cities, the estimated exponent at the end of the period is close to one.

We also estimate the Gini coefficients for each period, which have the advantage of not imposing a specific size distribution (Pareto for rank-size coefficients).[10] The results are similar: throughout the whole period, the evolution of the distribution indicates a divergent pattern as the coefficient rises from 0.45 in 1860 to 0.61 in 1960, with it growing particularly fast after 1930. Finally, Graph (c) in Figure 2 shows the empirical density functions for the four periods (estimated using adaptive kernels). It can be seen that the distribution remained almost static from 1860 to 1900. Since then, from a highly leptokurtic distribution with much of the density concentrated in the mean value of the distribution, it has lost kurtosis and the concentration has decreased. This evolution is more pronounced for the sample of 262 largest municipalities in 1860 (Graph (d)), indicating that the initially largest cities were especially involved in the divergence process. Thus, once more, our results point to increasing inequality within the distribution.

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[9] We also estimate the Pareto exponent using simple OLS regressions and the Hill estimator, and the results are quite similar.

[10] However, there is a statistical relationship between Zipf's law and the main concentration indices: Gini, Bonferroni, Amato and the Hirschman–Herfindahl index (Naldi, 2003).

Both analyses, the parametric and nonparametric one, show that the distribution remained stable until 1900, when a process of divergent growth started. Our results are robust to the entry of new cities in the sample, because when we consider a balanced panel of cities, we obtain similar patterns. The hypothesis we test in the following sections is that the factor driving this change in the distribution of city sizes is the economic integration process that took place during the period 1860 to 1960, the effects of which were particularly marked after 1900.

4. Data

To analyse the growth in Spanish cities, we use, as in the previous section, official city population data from the decennial censuses. Population data for 1860 are from Reher (1994), and for all the other decades, our data source is the census conducted by the National Statistics Institute (*Instituto Nacional de Estadística*) INE. Our main hypothesis is that the domestic market integration that took place between 1860 and 1960, under the presence of agglomeration economies, was a fundamental cause of the change in the structure of Spanish cities. Therefore, our main explanatory variable is market potential, which reflects the market access of each city. This variable has been extensively used in recent studies focusing on the determinants of growth and spatial distribution of cities, including Black and Henderson (2003), Ioannides and Overman (2004) and Bosker et al. (2008).

However, one of our empirical contributions is that we do not use the distance-weighted sum of the population of all other existing cities as a proxy of a city's market potential. Instead of this common option, we use the market potential variable from a retrospective estimate of regional market potential that is distributed across cities based on the relative size of the cities in each region. The regional market potential is the so-called 'nominal market potential' or the Harris (1954) market potential equation, defined as:

$$MP_i = \sum_{j-1}^{j-n} \frac{M_j}{d_{ij}} , \qquad (3)$$

where M_j is a measure of the size of province[11] j (GDP) and d_{ij} is the distance, or in this case, the bilateral transport costs between i and j.

By employing this expression, Martínez-Galarraga (2010) offers a measure of Spanish NUTS3 market potential for the years 1860, 1900 and 1930 based on Crafts' study (2005).[12] The author obtains historical market potential figures for Spanish NUTS3 regions as follows. First, he considers that market potential can be divided into two main components. Thus, he calculates domestic market potential (DMP_r) and foreign market potential (FMP_{rf}) between the provincial and international node *f*. In particular, the market potential of a province *r* (MP_r) is calculated as the sum of the domestic and foreign market potential: $MP_r = DMP_r + FMP_{rf}$.

Following this expression, the domestic market potential of each of the 47 provinces *r* is calculated as the sum of two components:

$$DMP_r = \sum_{1}^{s=46} \frac{M_s}{d_{rs}} + SP_r \,, \tag{4}$$

with $SP_r = \dfrac{M_r}{d_{rr}}$ the measure of the self-potential of each province *r*, where d_{rr} is calculated by taking a distance θ_{rr} equivalent to a third of the radius of a circle with an area equal to that of the province: $\theta_{rr} = 0.333\sqrt{\dfrac{(areaoftheprovince_r)}{\pi}}$.

The size of provincial markets (M_r) is measured in terms of aggregate income. GDP data at the NUTS3 levels are obtained from Rosés et al. (2010). The distance between regions d_{rs} is calculated including transport costs, which requires access to data on distances and average transport rates for commodities. Internal transport is assumed to be by railway and coastal shipping. For railway distances, the Ministry of Public Works (*Ministerio de Obras Públicas*) (1902) and Wais (1987) were consulted. For distances between ports, electronic atlases supply information on the length of sea journeys.[13] For transport costs, data on railway rates were obtained from Herranz (2006) and coastal shipping rates in 1865 were

[11] Provinces are Spanish NUTS3 regions.
[12] See Martínez-Galarraga (2010) for a detailed description.
[13] www.dataloy.com and www.distances.com.

obtained from Nadal (1975). In order to consider the reduction in sea transport costs, the data were corrected with the freight rate indices calculated by Mohammed and Williamson (2004). However, in 1860, our first benchmark year, only 32 of the 47 provinces considered were connected to the railway network. For this reason, road transport was also included in the domestic market potential estimates for this year.[14] Distances by road were taken from the General Directorate of Public Works (*Dirección General de Obras Públicas)* (1861). For road transport prices, the information provided by Barquín (1999) was used. Finally, the relative weight of each transport mode in the coastal provinces was obtained from Frax (1981).

In Martinez-Galarraga (2010), foreign markets were added to domestic market potential as follows. The construction of foreign market potential (FMP_{rf}) is based on the gravity equation for international trade estimated by Estevadeordal et al. (2003). The elasticities obtained for distance and tariffs are used here to reduce the size of foreign markets. The selection of foreign markets is based on the geographical distribution of Spanish exports, which reveals a high concentration in export markets (Prados de la Escosura, 2000; Tena, 2005). On the basis of this information, countries that accounted for at least 5% of Spain's exports are selected as foreign markets.[15] Thus, four countries are considered in the calculation of foreign market potential: Great Britain, France, Germany and the United States.[16] Having decided on which countries to include in the sample, the next step involves selecting a node to calculate the distance from Spanish provinces to each of the four markets. In the case of Great Britain, London – the capital and economic centre of the country – is selected.[17] For the US, the choice is New York, while in the case of Germany, for questions of geographical access and the size of its port, the city of Hamburg is taken as the node. However, in the case of France, the way of proceeding must differ. As a consequence of its geographical location in relation to that of the Iberian Peninsula, the

[14] In 1930, however, road transport was not yet playing an important role, and therefore, it was not considered (Herranz, 2006).

[15] Two exceptions include Cuba, a market that received a high percentage of Spanish exports, above all in the mid-19th century (18.5% of the total but only 5.3% in 1913 and 2.1% in 1930), and Argentina, whose market exceeded the 5% threshold on the eve of the First World War. They are excluded due to data restrictions regarding GDP at current prices. However, it ought to be the case that the limited size of their markets and, especially, the great distance separating them from the Peninsula would minimise the cost of their exclusion.

[16] Overall, these four countries accounted for 62.4% of Spanish exports in 1865/69, 57.8% in 1895/99, 58.0% in 1910/13 and 58.9% in 1931/35, with France and Great Britain the main markets.

[17] Crafts (2005) gives disaggregated information for regional GDP in Great Britain. Hence, it is possible to calculate market potential, not by assigning all the economic activity in Britain to London but rather by distributing it between the nodes selected for each of the 12 regions. However, this approach provides similar results.

French market can be accessed both via the Atlantic and via the Mediterranean seaboards. Therefore, localising the French market into a single node would mean penalising the regions on one or other of these two seaboards. For this reason, the French market is divided to capture the various routes along which Spanish provinces can access it. Thus, three regional nodes are considered: Le Havre and Nantes on the Atlantic seaboard and Marseille on the Mediterranean. The GDP of the main trading partners of Spain was obtained from Crafts (2005) based on the estimates of Prados de la Escosura (2000). Prevailing exchange rates were applied to convert the GDP figures from pounds to pesetas. Maritime distances were once again obtained from an electronic atlas and tariffs from O'Rourke (2000) and Mitchell (1998a, 1998b).

The foreign market potential of province r $\left(FMP_{rf}\right)$ is thus obtained according to the next expression, where d_{rp} captures the distance from the inland provincial node to the nearest Spanish port: $FMP_{rf} = \sum_{1}^{f=4} \frac{M_f}{d_{rp}} \cdot Distance^{-0.8} \cdot Tariff^{-1.0}$, with $d_{rp} = 1$ if r is a coastal province and $d_{rp} = d_{rs}$ if r is an inland province.

Hence, the market potential of province r $\left(MP_r\right)$ is obtained as the sum of the following terms, the first two corresponding to domestic market potential (including the self-potential of province *r*) and the last one capturing foreign market potential:

$$MP_r = \sum_{1}^{s=46} \frac{M_r}{d_{rs}} + SP_r + \left[\sum_{1}^{f=4} \frac{M_f}{d_{rp}} \cdot Distance^{-0.8} \cdot Tariff^{-1.0} \right], \qquad (5)$$

with d_{rp} conditioned to the coastal or inland nature of province *r*.

By departing from the different components of provincial market potential constructed by Martinez-Galarraga (2010) and described above, we calculate the market potential for each city. Thus, we define city market potential as the sum of three elements:

$$MP_{i,r} = \frac{Pop_i}{\sum_{i=1}^{n_r} Pop_i} \cdot DMP_r + \sum_{j \neq i}^{n_r - 1} \left(\frac{1}{d_{ij}} \cdot \frac{Pop_j}{\sum_{j \neq i}^{n_r} Pop_j} \cdot DMP_r \right) + FMP_r . \qquad (6)$$

The market potential of city *i* in region *r* is the sum of (1) the estimated own market potential of city *i*, calculated from province domestic market potential weighted by the relative size of the city (measured by the share of total provincial population),[18] (2) the sum of the domestic market potentials of all the other $j \neq i$ cities within the region weighted by the inverse physical distances, each calculated using their population relative sizes, and (3) province foreign market potential.

Finally, we also introduce several geographical variables into the estimations to control for *first nature* causes. Altitude and ruggedness data by municipality were obtained from Azagra et al. (2006) and Goerlich and Cantarino (2010), respectively.

5. Empirical analysis

First, we conduct a nonparametric analysis of the effects of market potential on urban growth. To do this, we estimate the nonlinear relationship between initial market potential and growth using a local polynomial smoothing[19] for any cross-section in our sample. Figure 3 shows the results, including the 95% confidence bands. These graphs show the presence of heterogeneity over time in the impact of market potential on city growth. We can observe a marked temporal evolution pointing to the increasing influence of market potential over time for the cities with high initial market potential. In the 1860–1900 period, the relationship between mean population growth and initial market potential is similar for all the distribution of market potentials, although there are small fluctuations across the distribution. However, in the next two periods (1900 to 1930 and 1930 to 1960), the graphs display a clear evolution to a U-shaped effect of market potential on growth. Although from 1900 to 1930 the relationship is still almost a flat line for most of the market potentials, a positive relationship between initial market potential and growth appears for the cities with the highest initial market potentials. Finally, in the 1930 to 1960

[18] Therefore, the self-potential of each city is obtained from provincial market potential in proportion to the population of the city over the overall provincial population. For example, the population of Madrid city in 1860 was 57% of the total population of the Madrid province, and this is the same proportion that we apply to province domestic market potential to obtain the own market potential of Madrid city. This value can be considered to be a lower bound because the share of province market potential corresponding to the largest cities is probably greater than their proportion of the total population, as nonlinear behaviours and increasing returns to scale can be involved in a NEG framework.

[19] The local polynomial smoother fits the growth rate $g_{it} = \left(\ln S_{it+1} - \ln S_{it}\right)$ to a polynomial form of $\ln MP_{it-1}$ via locally weighted least squares. We use the lpolyci command in STATA with the following options: local mean smoothing, a Gaussian kernel function to calculate the locally weighted polynomial regression and a bandwidth determined using Silverman's (1986) rule of thumb.

period the U-shaped pattern has clearly emerged; at the beginning of the distribution the impact of market potential on growth decreases with market potential, but at high levels of initial market potential the effect is positive and increasing with market potential. This pattern points to growing inequality within the distribution of market potentials.

Herein Figure 3

Second, by conducting a parametric analysis, we want to exploit the panel structure of our data. Therefore, we study the period 1860–1960 using panel data and consider three homogeneous periods: 1860–1900, 1900–1930 and 1930–1960.[20] The initial and final sets of cities are those plotted in Figure 1. Our baseline equation is similar to that proposed by Black and Henderson (2003) and Henderson and Wang (2007):

$$g_{it} = \alpha + \beta \ln MP_{it-1} + \phi g_{it-1} + \delta X_{it} + \varepsilon_{it}, \qquad (7)$$

where the independent variable is the logarithmic growth rate, $g_{it} = (\ln S_{it} - \ln S_{it-1})$. The main explanatory variable is city market potential (MP_{it}) and X_{it} is a vector of both the time-variant and the time-invariant variables.

The time-invariant variables represent the locational fundamentals (*first nature* factors) of each location. They include a coastal dummy indicating if the city has access to the sea, the city's altitude and a measure of ruggedness taken at the city level. A dummy for each province's capital is included, as city growth may have been affected by the administrative status of the city, as well as a Madrid dummy (equal to one only for the city of Madrid and zero otherwise) to account for the specific agglomeration in the capital of the country. The Madrid dummy reflects that the capital city tends to be more dominant the more political instability there is in a country and the more authoritarian is its regime (Ades and Glaeser, 1995), given that Spain suffered military dictatorships during the 20th century (1923–1931 and 1939–1975). Ayuda et al. (2010) also introduce this Madrid dummy variable and find it to be significant. Regional fixed-effects (NUTS2 regions) are also included to control for other local characteristics for which we have no data. The log–log specification simplifies the interpretation of coefficients (elasticities). A positive β coefficient is expected, as city growth increases in those locations with greater market potential. Although this positive

[20] We must consider homogeneous periods because, if we mixed different time periods (1860–1900 and decennial data since 1900) the later periods could overpower the estimated effects just because we would have more decade observations.

effect is expected to fall with the market potential size for many cities, given the U-shaped behaviour previously observed in most of the periods. This pattern is especially clear for the last period (1930–1960). Finally, we add the lagged growth rate to the specification[21] to control for the persistence in growth rates.

Herein Table 2

Table 2 shows the correlations between the variables. The first column reports the correlations between the explanatory variables and our endogenous variable, city growth. City market potential is significantly correlated with city population growth with a positive correlation. All the other variables have significant correlations with growth. The low correlation between population growth and past growth (-0.069) is surprising because other studies usually find a high persistence in growth rates (see Glaeser and Shapiro (2003) for the US case). The explanation is that we consider time periods higher than one decade (1860–1900, 1900–1930 and 1930–1960), so deviations in growth paths are more possible, especially in a historical context of internal market integration. It is noteworthy that the correlation between growth and provincial market potential is the lowest one (0.041) and significant only at the 10% level. Our explanation for this low correlation is that, although provincial market potential is the same for all the cities within the region, there is high variation in growth rates across cities, and thus provincial market potential does not help to explain that variation in growth rates.[22]

Herein Table 3

Table 3 shows the OLS estimates of Eq. (7). We run the regression for each sub-period and for a pool 1860–1960 including all the observations. Columns 1 a 2 display the estimates for the 1860-1900 period. Surprisingly, the association between city growth and market potential is not significant with (column 2) and without control variables (column 1). However, as we consider more periods we obtain a positive and significant effect of market potential in 1900-1930 (column 4), 1930-1960 (columns 5 and 6) and the pool 1860–1960 including all the observations (columns 7 and 8). Moreover, the impact of market potential

[21] Note that as new cities enter the sample over time and Eq. (7) includes past population growth rates, the total number of observations in our regressions does not coincide with the sum of the number of cities by period shown in Table 1.
[22] Nevertheless, we also estimate using provincial market potential instead of city market potential, and the results are similar. These results are shown in the Appendix, Tables A1 (OLS results) and A2 (IV results).

clearly increases over time, from a non-significant -0.026 in 1860-1900 to a positive and significant in 0.052 in 1930-1960. The coefficients in 1900-1930 and 1930-1960 are almost equal. The estimated effect for the pool 1860-1960, 0.043, is an average value.[23] It is interesting that in the first period (1860-1900) market potential is not significant but the Madrid dummy is strongly positive while simultaneously past growth exerts a significantly negative impact. This period seems to experience a bi-polarisation of Spanish cities, with Madrid growing a lot while small cities were catching-up with respect to other larger cities. This contrast with the next two periods where full divergence is observed: not only market potential matters but the effect of past growth, which changes to non-significant in 1900-1930 and finally strongly positive in 1930-1960, and Madrid still growing faster in the 1900-1930 period.

Remember that Figure 3 showed a decreasing effect of market potential on growth for the smallest values of market potential in some periods. Therefore, to model the possible heterogeneous effects of market potential on growth, we also estimate quantile regressions (Koenker and Bassett, 1978). The quantile regression version of the linear model shown in Equation (7) can be written as

$$g_{it} = \alpha(\tau) + \beta(\tau)\ln MP_{it-1} + \phi(\tau)g_{it-1} + \delta(\tau)X_{it} + \zeta_{it}. \qquad (8)$$

Note that the estimated parameters are τ-dependent in this case, where τ is the corresponding quantile of the growth rate. Thus, quantile regressions provide a richer characterization of the data, allowing us to consider the impact of the market potential on the entire distribution of g and not merely its conditional mean. Quantile regressions take into account unobserved heterogeneity and allow for heteroskedasticity among the disturbances, non-normal errors, and are more robust to outliers than standard OLS regressions.

Figure 4 shows the quantile regression results for the model of Equation 8. The graphs display the estimates of the coefficient and the 95% confidence intervals for the market potential across the nine quantiles considered (ranges from 0.1 to 0.9). We consider each sub-period and the pool 1860–1960, and we estimate with and without control variables.

[23] This positive effect is expected to fall with the market potential size for many cities, given the transition to a U-shaped pattern previously observed in Figure 3. To check this possible nonlinear relationship, we re-run Eq. (7) using the squared market potential variable, $\gamma(\ln MP_{it-1})^2$, instead of $\beta \ln MP_{it-1}$, expecting a γ positive coefficient. The results reveal a slight nonlinear relationship (estimated coefficients are between 0.003 and 0.004) only significant in the 1930-1960 period (as Figure 3 shows) and for the pool 1860–1960. These results are available from the authors on request.

Thus, each graph shows the nonlinear counterpart of the linear estimates of the market potential coefficient previously shown in Table 3. When the only explicative variable is the market potential (graphs 1, 3, 5 and 7), we find a clear nonlinear effect of market potential, although in some sub-periods (1860–1900 and 1930–1960) these differences across quantiles are not significant. This means that, the higher the quantile, the greater the impact of market potential on growth. In particular, the quantile estimates show that for the pool 1860–1960 (graph 7) the effect of market potential is more than 4 times higher in the top quantile (0.9) than in the bottom quantile (0.1), as the coefficient increases from 0.011 to 0.047, pointing to a higher effect of market potential on growth in those cities with higher growth rates. However, when we add all the controls (graphs 2, 4, 6 and 8) the nonlinear relationship eventually disappears and coefficients across quantiles are similar across quantiles in all periods. Thus, this indicates that linear estimates of Equation (7) are robust across the distribution of cities if all controls are added.

Nevertheless, there can be an issue of endogeneity in the specification of the model of Equation 7. As Henderson and Wang (2007) highlight, unobservable effects related to individual cities may operate at a regional level and may be correlated with city growth and market potential: some geographical characteristics, local culture, business climate or institutions. Furthermore, in any city growth estimation, there could be an issue with the time persistence in the error structure. Moreover, there might be potential spatial correlation between cities. Because of that, we consider that infrastructures are a key element to explain changes in market potential for cities, while another concern relates to the role played by these infrastructures. Policymakers tend to improve infrastructures in the most populated cities, but these infrastructures (roads, railways, etc.) undoubtedly also increase the market access of these locations (Holl, 2012; Garcia-López et al., 2015). The policy decision process and construction of these infrastructures often take several periods, so the past growth rate we introduce in the specification can incorporate the forecasting of these infrastructures, and hence this dynamic model could alleviate the possible endogeneity problem.

We need to instrument the market potential variable in the first-stage regressions of the IV estimation. We use two instruments: a measure of distances across cities and the lagged regional population density.[24] The measure of distances across cities is defined as $\sum_{j\neq i}^{n_t-1}\frac{1}{d_{ij}}\Big/\frac{1}{n_t}\sum_{i}^{n_t}\sum_{j\neq i}^{n_t-1}\frac{1}{d_{ij}}$. Head and Mayer (2006) use $\sum_{j\neq i}^{n_t-1}\frac{1}{d_{ij}}$ as a measure of location *i* centrality to instrument regional market potential; we weight their centrality index by the average centrality across cities to use a relative measure of distances. Thus, this is a time invariant measure of the geographical location of cities. An alternative instrument would be to consider the geographical distance to the nearest central place. However, the choice of the reference points could raise another endogeneity issue because the central places usually are cities with high market potential. Our choice of a centrality index is a more flexible measure because it does not explicitly impose a centre.

On the other hand, population can serve as a good measure of market potential, and in some papers it is used directly instead of GDP (Black and Henderson, 2003; Ioannides and Overman, 2004; Henderson and Wang, 2007). To be cautious, we use the lagged values of population density.[25] It could be argued that using regional population density might not be a good instrument, because it is probably highly correlated with city growth, our endogenous variable, in the second-stage regression. However, one of the particular features of the Spanish urban system was that internal migrations were not statistically related to differential city growth for the whole sample of cities in 1930. Silvestre (2005) shows that although Spanish internal migrations grew significantly during the 1920s, these movements were limited to just a few big cities. In fact, Silvestre calculates that by 1930, two provinces, Madrid and Barcelona, accounted for 45.97 % of the total stock of Spanish migrants. Figure 5 confirms this idea. It shows the scatterplot of city growth against the lagged regional population density for our panel from 1860 to 1960. The graph includes all observations, and no clear pattern of any kind is found. The fitted line is almost a horizontal line, and the slope (0.0004) is not significant. Therefore, we expect that neither the lagged regional population density nor the centrality index (which represents the exogenous geographical location of cities) have a direct causal effect on our endogenous variable (city population growth), and thus our instruments satisfy the exclusion restriction.

Herein Figure 4

[24] We use regional population density instead of city population density because provincial boundaries remain constant over time, while city boundaries change in some cases.

[25] Thus, population density values from 1787 are used to estimate market potential in 1860, 1860 values to estimate market potential in 1900 and the values from 1900 to estimate market potential in 1930.

Table 4 reports the IV results. The Madrid dummy cannot be included in any of the cross-sectional estimations; this is a singleton dummy (a variable with a value of 1 and a 0 value for all the other observations), and it causes the robust covariance matrix estimator to be less than full rank. Thus, we can only include it in the regression for the pool 1860-1960 (column 8). We estimate the second-stage regressions using GMM; some statistics from the first-stage regressions are also reported.[26] Overall, our instruments seem to perform well, as the weak instrument hypothesis is rejected using the Stock–Yogo test and the model passes the underidentification test at any significance level. The null of the overidentification test (Hansen J statistic) cannot be rejected in most of the cases, although models in column 1, 5, 6 and 7 do not pass the test at the 5% conventional level. We obtain a negative coefficient for the effect of market potential on growth in the 1860-1900, but the coefficients are non-significant (columns 1 and 2). From the 1900-1930 period the coefficient is significant and positive when all controls are added (columns 4 and 6). This means that the significant effect of market potential estimated for the pool 1860-1960 is basically driven by the pattern from 1900, when the U-shaped pattern emerged, confirming the existence of increasing agglomeration economies over time. The estimated average value for the pool 1860-1960, 0.104 (column 8), is higher than that obtained by OLS, and it is closer to the raw correlation shown in Table 2. This estimated elasticity means that a 1% increase in market potential implies an average 0.10% increase in the population of the city. Other authors also find a heterogeneous effect of market potential over time. Combes et al. (2011) use different measures of population, employment and value-added for each French department from 1860 to 2000. They find that the French economy has exhibited agglomeration economies, which seem to have reinforced over time. However, the spatial distribution of these gains is determined mainly by market potential in the first sub-period, 1860–1930, but in the second period (1930–2000) human capital is the main determinant.

Herein Table 4

Finally, we adopt an alternative perspective and try another definition of city market potential. The strength of our results comes from the quality of our data, because we use market potential constructed using a set of historical estimates of regional GDP and historical transport costs among regions. However, a common approach in the literature

[26] The complete results of the reduced regressions, first-stage regressions and all the tests, not shown for size restrictions, are available from the authors on request.

consists of taking a distance-weighted sum of the population of all other existing cities as a proxy of city market potential (Black and Henderson, 2003; Ioannides and Overman, 2004; Bosker et al., 2008). Thus, following Black and Henderson (2003), we define market potential as

$$mp_{it} = \sum_{i \neq j} \frac{Pop_{it}}{d_{it}}. \qquad (9)$$

Market potential is the sum of the populations (Pop_{it}) of all cities weighted by physical distances (d_{it}), calculated using the geographical coordinates. We re-estimate Eq. (7) considering this market potential definition based on population to be our main explanatory variable and using the same set of controls and instruments as above. Table 5 shows the OLS results.

Herein Table 5

These results are similar to those obtained by using the market potential calculated from regional GDPs (Table 3), although the estimated coefficients are lower. Thus, in the first sub-period (1860-1900) the relationship between growth and market potential is not significant, but in the rest of sub-periods and the pool 1860–1960 we obtain a positive and significant effect of market potential.

Herein Table 6

Finally, Table 6 shows the IV estimates. Again, the results are similar to those obtained in Table 4, finding a positive effect of market potential on city growth only significant for the period (1900-1930), which may be driven the estimated effect for the pool 1860-1960. Nevertheless, these results are less robust; our instruments may perform worse with market potential based on population. The model does not pass the overidentification test (Hansen J statistic) in the last period (1930-1960) and the pool 1860-1960, and the underidentification and the weak identification tests are not passed in columns 1 and 2.

6. Conclusions

In this paper, we analysed the role of the increase in market potential on city growth, focusing primarily on the factors that determine the configuration of the urban system.

Our hypothesis was that in the context of the growing economic integration and industrialisation experienced by Spain from the second half of the 19th century onwards, the presence of agglomeration economies implied that access to markets was an important factor explaining the relative growth of Spanish cities and thus affecting the country's urban pattern. In order to test this hypothesis, in the empirical analysis both location fundamentals (relating to Spain's geography) and city market size (a variable introduced by theoretical approaches conducted within NEG) were considered to be determinants of city growth. Our results show that Spain's urban growth was related initially to *first nature* characteristics but that it was also affected by the forces of agglomeration economies, measured in terms of home market size. All these elements had an impact on Spanish urban design, which experienced a growth in its city size inequality because throughout the process of economic growth and integration, the cities with greatest market potential benefited most from the presence of agglomeration economies.

The present analysis confirms the results obtained in works devoted to the study of the determinants of population growth in US, Brazilian or Chinese cities during different periods in the second half of the 20th century. In this respect, our main contribution is the use of a direct measure of market access. This was constructed by making use of regional GDP and transport costs at the time of each of the benchmark years. The results point out the presence of a positive effect of market potential on population growth in Spanish cities.

Our results also fall in line with those that have pointed to the importance of agglomeration forces in the definition of Spanish economic geography. A bulk of the NEG empirical literature devoted to the analysis of the Spanish historical experience has shown that market access acts as a relevant factor explaining regional industry location, the upsurge of wage gradients and the direction and intensity of internal migratory flows along the process of national market integration and industrialisation. In this paper, it was shown that the market potential of cities was also a key factor in explaining their relative growth during the period 1860–1960, especially from the very beginning of the 20th century.

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Table 1. Number of cities and descriptive statistics by year

Year	Cities	Mean population	Standard deviation	Minimum	Maximum
1860	266	13,267.34	24,712.85	5,004	279,379
1900	657	13,720.86	32,795.34	5,001	539,835
1930	877	15,684.87	50,554.06	5,000	1,005,565
1960	1,030	20,878.10	91,372.98	5,004	2,259,931

Data sources: Reher (1994) and *Instituto Nacional de Estadística*, www.ine.es.

Table 2. Raw correlations

	Population growth	City Market potential	Provincial Market potential	Population growth (t-1)	Altitude	Ruggedness	Coastal dummy	Madrid dummy	Capital dummy
Population growth	1								
City Market potential	0.193***	1							
Provincial Market potential	0.041*	0.776***	1						
Population growth (t-1)	-0.069**	0.047	-0.056*	1					
Altitude	-0.176***	-0.316***	-0.337***	-0.176***	1				
Ruggedness	-0.043**	0.064***	0.096***	-0.043**	0.022*	1			
Coastal dummy	0.131***	0.122***	0.185***	0.131***	-0.478***	0.111***	1		
Madrid dummy	0.056***	-0.015	0.003	0.056***	0.033***	-0.025**	-0.017	1	
Capital dummy	0.188***	-0.111***	-0.040**	0.188***	0.019	-0.083***	0.048***	0.139***	1

Notes: Pearson correlations. Correlations statistically different from zero at the *10%, **5% and ***1% levels.

Table 3. Spanish city size growth: City Market potential, OLS estimates

	1860-1900		1900-1930		1930-1960		Pool 1860-1960	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ln(Market potential)	0.023	-0.026	-0.002	0.051*	0.056***	0.052***	0.031***	0.043***
	(0.015)	(0.034)	(0.011)	(0.026)	(0.014)	(0.016)	(0.008)	(0.015)
Population growth (t-1)		-0.358***		0.025		0.297***		-0.031
		(0.094)		(0.062)		(0.053)		(0.047)
Madrid dummy		0.286***		0.293***		0.166		0.195
		(0.054)		(0.042)		(0.161)		(0.141)
ln(Altitude)		-0.004		-0.023		-0.042***		-0.037***
		(0.024)		(0.018)		(0.011)		(0.011)
ln(Ruggedness)		0.106**		-0.019		-0.017		0.001
		(0.043)		(0.025)		(0.018)		(0.015)
Coastal dummy		0.158**		-0.151**		-0.033		-0.046
		(0.077)		(0.064)		(0.032)		(0.033)
Capital dummy		0.213***		0.162***		0.295***		0.259***
		(0.059)		(0.038)		(0.038)		(0.031)
Regional fixed-effects	No	Yes	No	Yes	No	Yes	No	Yes
Time fixed-effects	No	No	No	No	No	No	Yes	Yes
Observations	255	166	608	255	797	589	1,660	1,010
R^2	0.007	0.541	0.001	0.269	0.041	0.437	0.070	0.254

Notes: 1) Market potential by city. 2) All the models include a constant. 3) Coefficient (robust standard errors). Significant at the *10%, **5%, ***1% level.

Table 4. Spanish city size growth: City Market potential, IV estimates

	1860-1900		1900-1930		1930-1960		Pool 1860-1960	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ln(Market potential)	0.029	-0.019	-0.048***	0.131***	0.091***	0.096***	0.028***	0.104***
	(0.019)	(0.052)	(0.016)	(0.041)	(0.014)	(0.023)	(0.010)	(0.023)
Population growth (t-1)		-0.340***		0.036		0.277***		-0.042
		(0.089)		(0.059)		(0.053)		(0.047)
Madrid dummy								0.224
								(0.138)
ln(Altitude)		-0.005		-0.014		-0.038***		-0.029***
		(0.024)		(0.016)		(0.011)		(0.011)
ln(Ruggedness)		0.115***		-0.005		-0.008		0.014
		(0.043)		(0.025)		(0.017)		(0.015)
Coastal dummy		0.169**		-0.160**		-0.039		-0.047
		(0.070)		(0.063)		(0.031)		(0.033)
Capital dummy		0.226***		0.207***		0.325***		0.289***
		(0.061)		(0.037)		(0.039)		(0.032)
Regional fixed-effects	No	Yes	No	Yes	No	Yes	No	Yes
Time fixed-effects	No	No	No	No	No	No	Yes	Yes
Underidentification test (Kleibergen-Paap rk LM statistic), p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Weak identification test (Cragg-Donald Wald F statistic)	204.434	35.341	241.600	43.305	622.343	199.956	829.819	229.701
Hansen J statistic, p-value	0.013	0.575	0.051	0.186	0.001	0.007	0.001	0.487
Observations	255	166	608	255	797	589	1,660	1,010

Notes: 1) Market potential by city. 2) Second stage regressions. 3) All the models include a constant. 4) Coefficient (robust standard errors). 5) Instruments: lagged provincial population density and a measure of geographical distances. Significant at the *10%, **5%, ***1% level.

Table 5. Spanish city size growth: City Market potential based on population, OLS estimates

	1860-1900		1900-1930		1930-1960		Pool 1860-1960	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ln(Market potential)	0.001	0.001	0.007***	0.015***	0.013***	0.005**	0.007***	0.005***
	(0.002)	(0.002)	(0.002)	(0.005)	(0.002)	(0.003)	(0.001)	(0.002)
Population growth (t-1)		-0.362***		0.005		0.289***		-0.035
		(0.094)		(0.063)		(0.057)		(0.048)
Madrid dummy		0.297***		0.370***		0.198		0.229
		(0.058)		(0.052)		(0.162)		(0.142)
ln(Altitude)		0.001		-0.028		-0.051***		-0.042***
		(0.025)		(0.018)		(0.010)		(0.010)
ln(Ruggedness)		0.114***		-0.026		-0.022		-0.006
		(0.042)		(0.025)		(0.018)		(0.015)
Coastal dummy		0.155**		-0.100		-0.020		-0.032
		(0.078)		(0.066)		(0.032)		(0.034)
Capital dummy		0.226***		0.162***		0.279***		0.244***
		(0.052)		(0.037)		(0.038)		(0.030)
Regional fixed-effects	No	Yes	No	Yes	No	Yes	No	Yes
Time fixed-effects	No	No	No	No	No	No	Yes	Yes
Observations	255	166	608	255	797	589	1,660	1,010
R^2	0.001	0.539	0.016	0.292	0.085	0.432	0.085	0.255

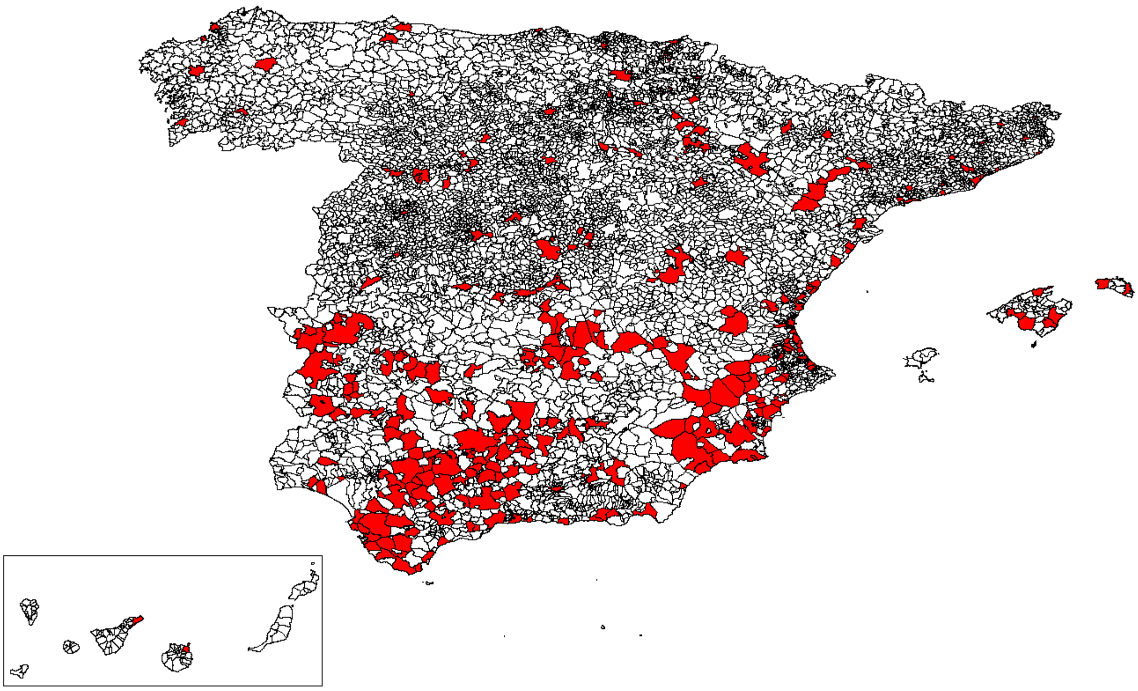
Notes: 1)Market potential by city based on populations. 2) All the models include a constant. 3) Coefficient (robust standard errors). Significant at the *10%, **5%, ***1% level.

Table 6. Spanish city size growth: City Market potential based on population, IV estimates

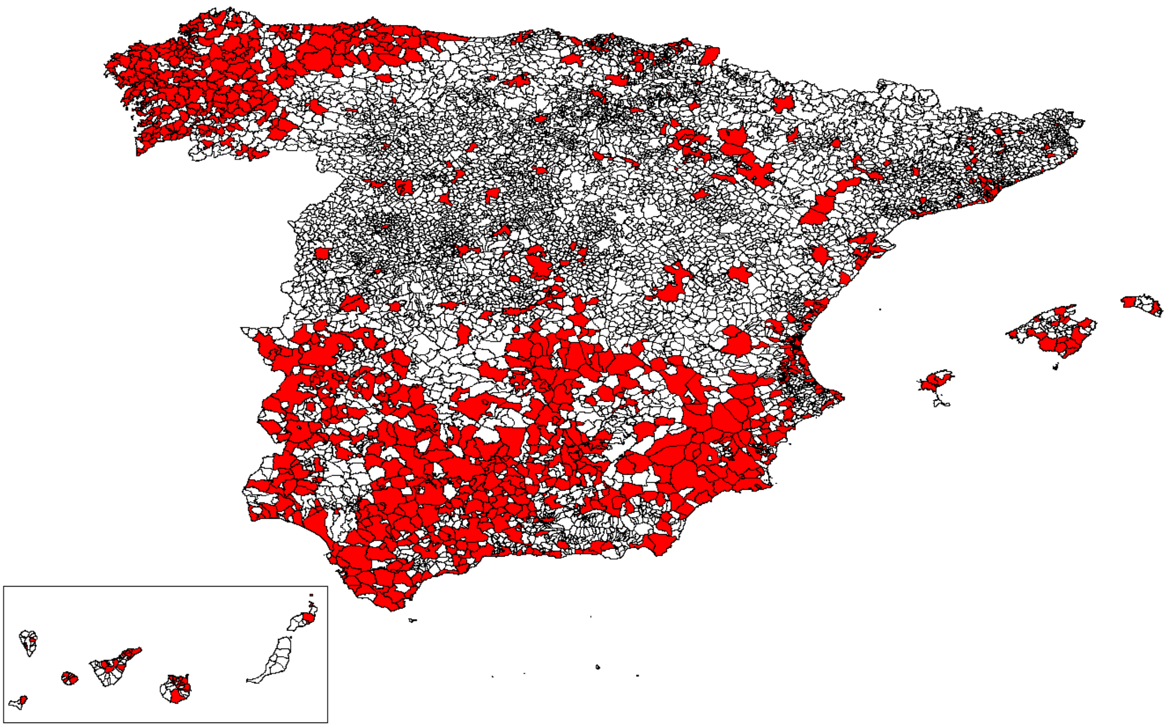
	1860-1900		1900-1930		1930-1960		Pool 1860-1960	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ln(Market potential)	-1.395	0.584	-0.674***	0.573***	-0.265**	0.073	-0.477***	0.346***
	(1.107)	(1.264)	(0.234)	(0.167)	(0.133)	(0.100)	(0.134)	(0.125)
Population growth (t-1)		-0.266		0.003		0.319***		-0.043
		(0.206)		(0.060)		(0.056)		(0.047)
Madrid dummy								0.308**
								(0.144)
ln(Altitude)		0.047		-0.029*		-0.055***		-0.042***
		(0.113)		(0.017)		(0.010)		(0.010)
ln(Ruggedness)		0.104*		-0.024		-0.005		0.003
		(0.056)		(0.024)		(0.017)		(0.015)
Coastal dummy		0.199*		-0.085		-0.048		-0.015
		(0.119)		(0.061)		(0.032)		(0.035)
Capital dummy		0.261***		0.185***		0.281***		0.248***
		(0.082)		(0.035)		(0.037)		(0.031)
Regional fixed-effects	No	Yes	No	Yes	No	Yes	No	Yes
Time fixed-effects	No	No	No	No	No	No	Yes	Yes
Underidentification test (Kleibergen-Paap rk LM statistic), p-value	0.304	0.767	0.000	0.000	0.000	0.000	0.000	0.000
Weak identification test (Cragg-Donald Wald F statistic)	1.654	0.236	45.110	65.088	78.769	231.144	63.185	30.019
Hansen J statistic, p-value	0.365	0.852	0.061	0.173	0.001	0.001	0.001	0.001
Observations	255	166	608	255	797	589	1,660	1,010

Notes: 1) Market potential by city based on populations. 2) Second stage regressions. 3) ll the models include a constant. 4) Coefficient (robust standard errors). 5) Instruments: lagged provincial population density and a measure of geographical distances. Significant at the *10%, **5%, ***1% level.

Figure 1. Cities in the sample, 1860–1900 and 1930–1960



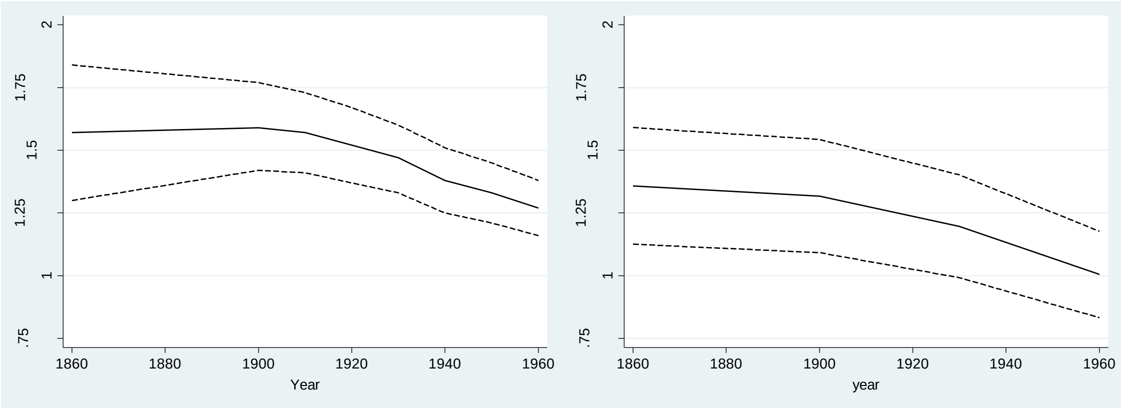
(a) 1860–1900



(b) 1930–1960

Notes: Geographical boundaries defined according to the census in 2000.

Figure 2. Evolution of the Spanish city size distribution



(a) Pareto: All sample (b) Pareto: Balanced panel (262 municipalities)

(c) Empirical pdfs: All sample municipalities) (d) Empirical pdfs: Balanced panel (262

Notes: The Pareto exponents are estimated by using Gabaix and Ibragimov's Rank-$1/2$ estimator. Dashed lines represent the standard errors calculated by applying Gabaix and Ioannides's (2004) corrected standard errors: $\text{GI s.e.} = \hat{a} \cdot (2/N)^{1/2}$, where N is the sample size.

Figure 3. Growth (ln scale) by initial market potential

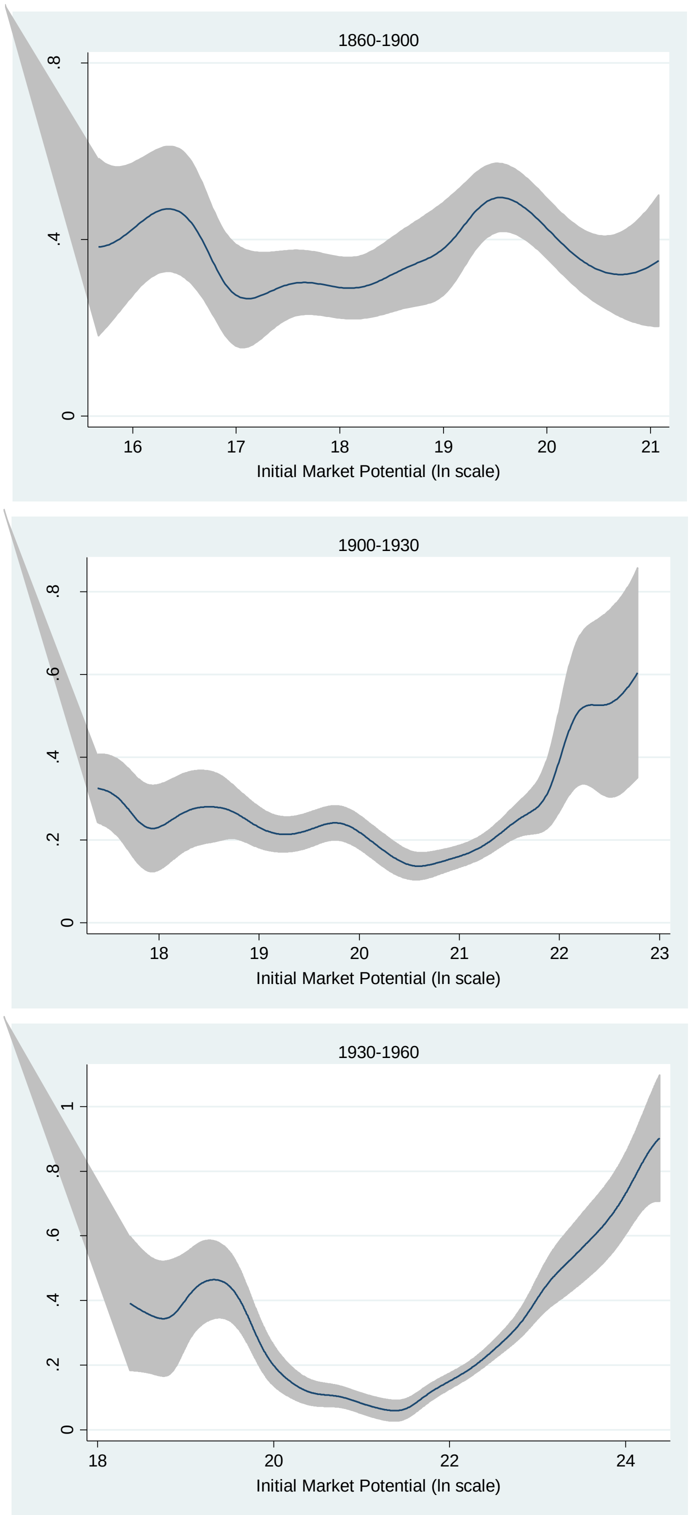
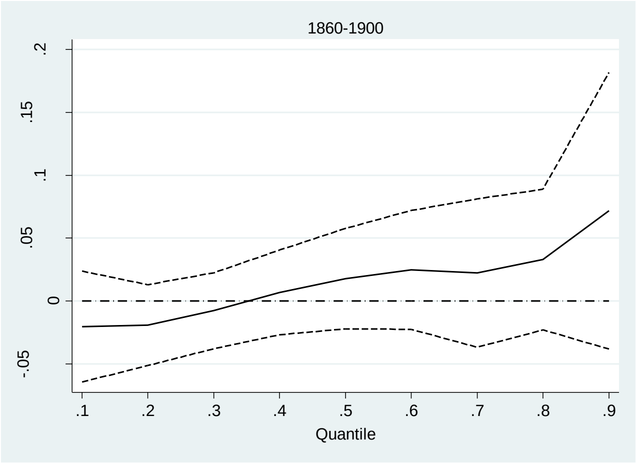
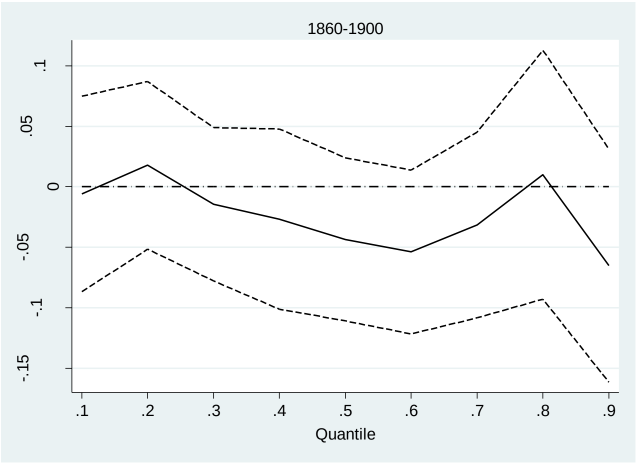


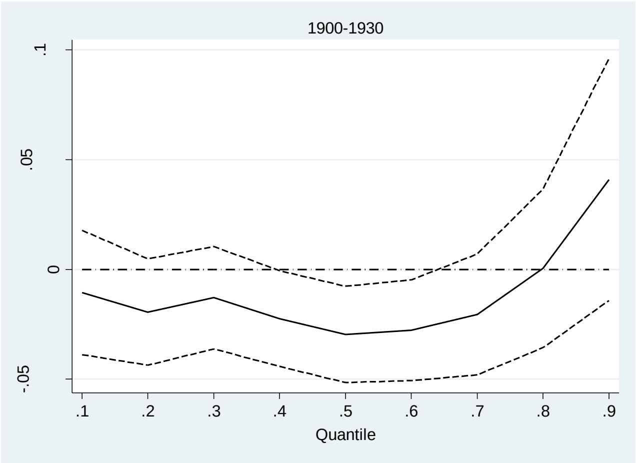
Figure 4. Quantile regression estimates: Growth vs. market potential

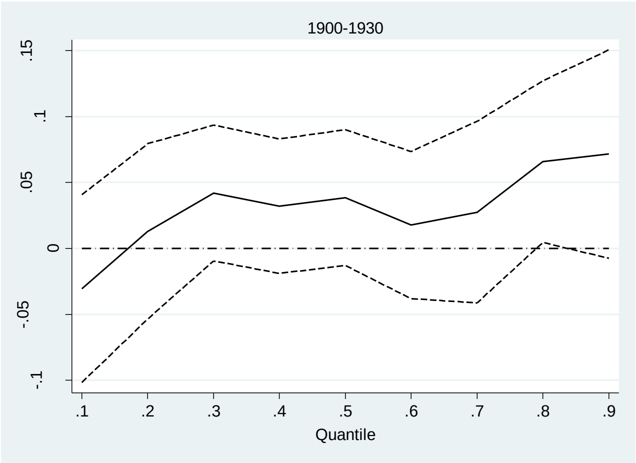




(1) 1860-1900: No controls

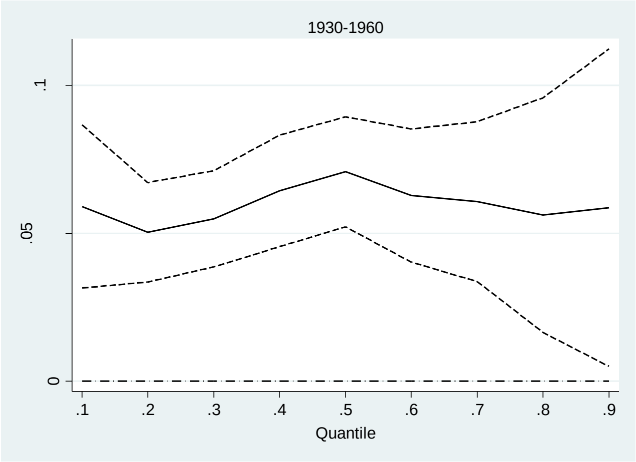
(2) 1860-1900: All controls

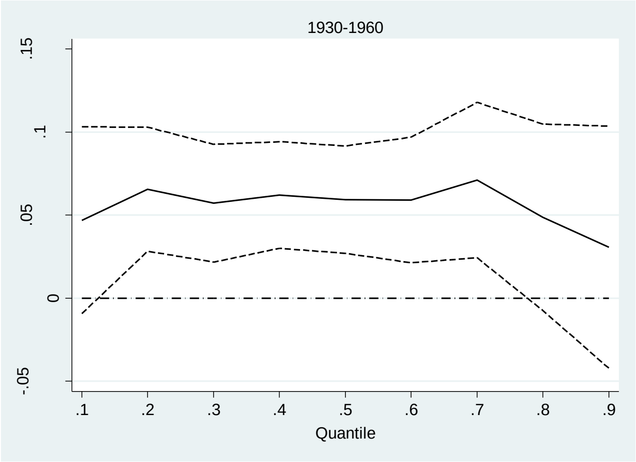




(3) 1900-1930: No controls

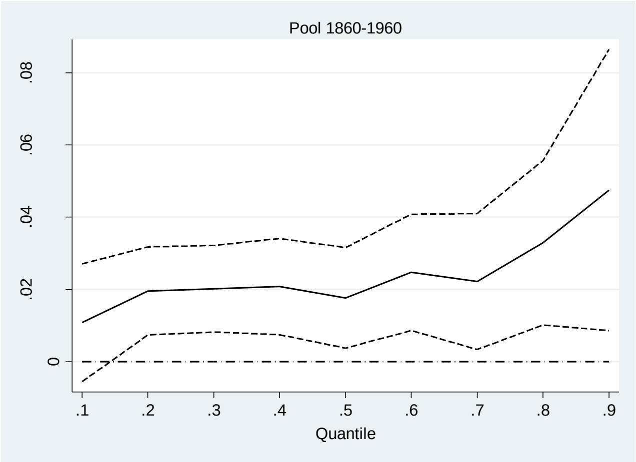
(4) 1900-1930: All controls

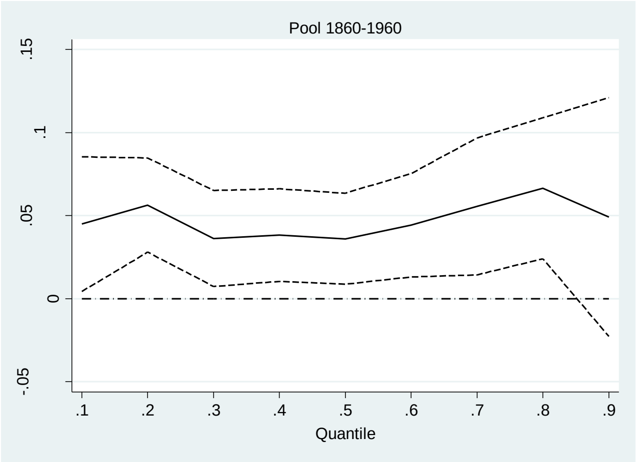




(5) 1930-1960: No controls

(6) 1930-1960: All controls



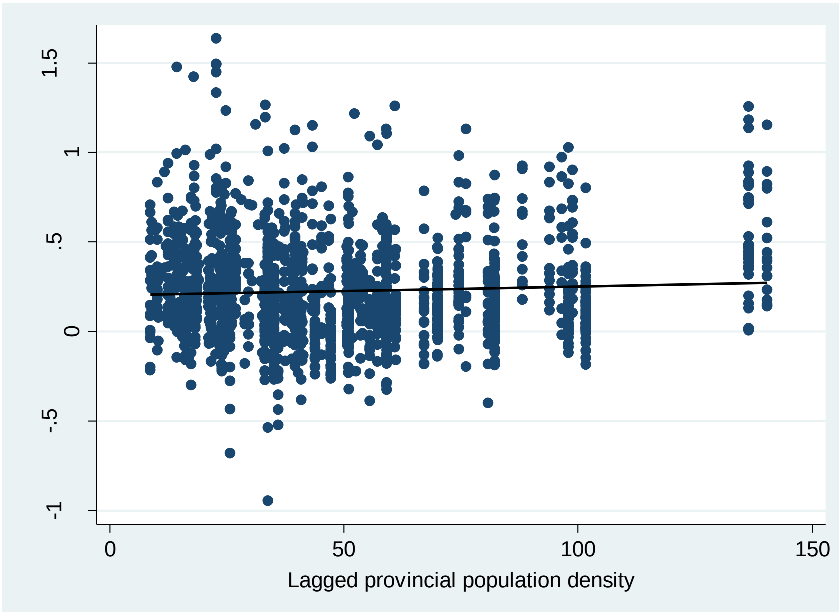


(7) Pool 1860-1960: No controls

(8) Pool 1860-1960: All controls

Note: Endogenous variable: logarithmic population growth rate. Models (1), (3) and (5) include a constant and the market potential (ln scale), and model (7) also includes time fixed-effects. In models (2), (4) and (7) all the controls are added: past population growth, a Madrid dummy, a coastal dummy, a capital dummy, altitude (ln scale), ruggedness (ln scale) and regional fixed-effects, and model (8) also includes time fixed effects.

Figure 5. Scatterplot of city growth (ln scale) against lagged provincial population density, Panel 1860–1960



Note: The linear trend is shown, although the estimated slope is not significant at the 5% level.

Appendix

Table A1. Spanish city size growth: Provincial Market potential, OLS estimates

	1860-1900		1900-1930		1930-1960		Pool 1860-1960	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ln(provincial Market potential)	0.059*	-0.102	-0.002	0.030	0.102***	0.068***	0.064***	0.042**
	(0.031)	(0.063)	(0.017)	(0.045)	(0.015)	(0.020)	(0.011)	(0.021)
Population growth (t-1)		-0.360***		0.027		0.314***		-0.023
		(0.090)		(0.063)		(0.058)		(0.048)
Madrid dummy		0.316***		0.269***		0.129		0.171
		(0.055)		(0.040)		(0.159)		(0.140)
ln(Altitude)		-0.012		-0.025		-0.040***		-0.037***
		(0.025)		(0.019)		(0.011)		(0.011)
ln(Ruggedness)		0.101**		-0.029		-0.024		-0.006
		(0.041)		(0.025)		(0.018)		(0.015)
Coastal dummy		0.166**		-0.147**		-0.036		-0.047
		(0.077)		(0.065)		(0.032)		(0.034)
Capital dummy		0.207***		0.143***		0.280***		0.244***
		(0.056)		(0.038)		(0.038)		(0.030)
Regional fixed-effects	No	Yes	No	Yes	No	Yes	No	Yes
Time fixed-effects	No	No	No	No	No	No	Yes	Yes
Observations	255	166	608	255	797	589	1,660	1,010
R^2	0.014	0.551	0.001	0.257	0.059	0.438	0.077	0.251

Notes: 1) Market potential by province. 2) All the models include a constant. 3) Coefficient (robust standard errors). Significant at the *10%, **5%, ***1% level.

Table A2. Spanish city size growth: Provincial Market potential, IV estimates

	1860-1900		1900-1930		1930-1960		Pool 1860-1960	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ln(provincial Market potential)	0.089**	-0.054	-0.044*	0.043	0.181***	0.130***	0.103***	0.120***
	(0.036)	(0.083)	(0.026)	(0.066)	(0.020)	(0.027)	(0.016)	(0.030)
Population growth (t-1)		-0.340***		0.039		0.310***		-0.012
		(0.086)		(0.060)		(0.051)		(0.047)
Madrid dummy								0.157
								(0.134)
ln(Altitude)		-0.010		-0.016		-0.031***		-0.028**
		(0.026)		(0.018)		(0.011)		(0.011)
ln(Ruggedness)		0.114***		-0.039*		-0.024		-0.003
		(0.040)		(0.023)		(0.017)		(0.015)
Coastal dummy		0.165**		-0.126**		-0.038		-0.053
		(0.071)		(0.062)		(0.031)		(0.033)
Capital dummy		0.224***		0.177***		0.295***		0.259***
		(0.053)		(0.035)		(0.038)		(0.030)
Regional fixed-effects	No	No	No	No	No	No	No	Yes
Time fixed-effects	No	Yes	No	Yes	No	Yes	Yes	Yes
Underidentification test (Kleibergen-Paap rk LM statistic), p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Weak identification test (Cragg-Donald Wald F statistic)	171.070	44.091	344.420	91.964	576.154	238.790	887.747	321.116
Hansen J statistic, p-value	0.090	0.899	0.002	0.001	0.001	0.591	0.001	0.071
Observations	255	166	608	255	797	589	1660	1010

Notes: 1) Market potential by province. 2) Second stage regressions. 3) All the models include a constant. 4) Coefficient (robust standard errors). 5) Instruments: lagged provincial population density and a measure of geographical distances. Significant at the *10%, **5%, ***1% level.