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LOCATION AND ENTREPRENEURSHIP: INSIGHTS FROM A SPATIALLY-EXPLICIT OCCUPATIONAL CHOICE MODEL WITH AN APPLICATION TO CHILE

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ABSTRACT. Occupational choice and heterogeneous managerial ability enter a spatial Dixit-Stiglitz setting, linking location, wages and regional entrepreneurship rates. Market potential has a positive partial effect and wages a negative partial effect on the regional supply of entrepreneurs, both balancing in equilibrium with endogenous wages. Market potential increases profits, but also the opportunity cost of entrepreneurship. In the long-run equilibrium with perfect mobility, the cut-off level of ability determining selection into entrepreneurship will be the same across regions; moreover, regional differences in entrepreneurship rates depend only in differences in average fixed costs of firms. An empirical application is provided for Chile.

1. INTRODUCTION

The relationship between economies of agglomeration and local entrepreneurship rates is not straightforward. Although usually documented to be positive (Audretsch and Fritsch, 1994; Van Soest et al., 2006; Sato et al., 2012), some empirical evidence suggests that the relationship could be insignificant or perhaps even negative (Naudé et al., 2008; Behrens and Robert-Nicoud, 2015). This paper presents a model that addresses, within a New Economic Geography (NEG) framework, the sometimes seemingly counter-intuitive relationship between the spatial distributions of entrepreneurship and regional market potential (as proposed by Hanson, 2005).

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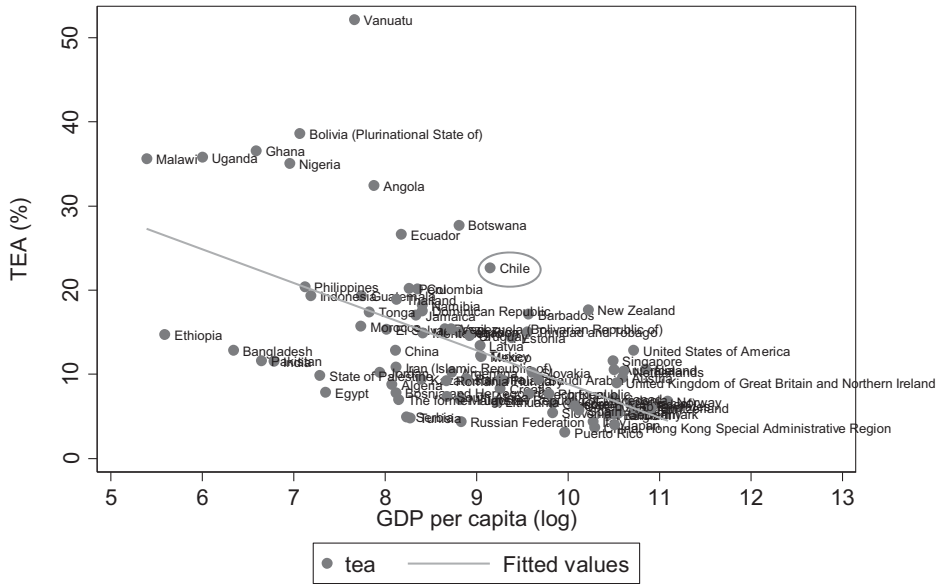
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Our model explicitly includes both push and pull mechanisms conditioning the spatial distribution of entrepreneurship. More specifically, we introduce occupational choice and heterogeneous managerial ability, two features of microeconomic theories of entrepreneurship (Parker, 2009), into the spatial Dixit-Stiglitz setting (Fujita et al., 1999). The resulting model predicts a positive partial effect of market potential, and a negative partial effect of domestic wages on the regional supply of entrepreneurs. But wages are also determined by the regional market potential, and so the model predicts, in the short-run equilibrium, offsetting effects of MP over the regional supply of entrepreneurs. The long-run equilibrium result, after full mobility of workers across regions and occupations, is that differences in entrepreneurship rates are given by differences in fixed costs across regions. We provide an empirical application of the model, by econometrically testing the offsetting effects of regional market potential on short-run changes in municipal entrepreneurship rates in Chile. The evidence supports the model's theoretical implications.

Our model introduces several innovations to existing NEG-based models of location and entrepreneurship. First, in contrast to Sato et al. (2012), we incorporate and test for interregional demand linkages as an explanation for the spatial variation in entrepreneurship, using a supply equation relating regional entrepreneurship rates and the Helpman–Hanson's market potential function (Helpman, 1998; Hanson, 2005). Second, in contrast to what are known as “footloose entrepreneurs” models of agglomeration (e.g., Forslid and Ottaviano, 2003), where there is an exogenous stock of entrepreneurs choosing their location, our model focuses on the effects of regional MP in the individual's decision to become an entrepreneur. Third, unlike Sato et al. (2012), our model endogenizes both business profits and wages, both conditioning the selection into entrepreneurship. The model yields testable parameter restrictions reflecting the trade-offs between regional incentives and the opportunity costs of entrepreneurship. Finally, it incorporates entrepreneurs' heterogeneity, leading to the sensible implication that, *ceteris paribus*, less efficient firms can succeed in areas of higher market potential and that average firm size should be greater in such places. Moreover, the result that long-run differences in local entrepreneurship rates are given by differences in fixed costs is consistent with location drivers in spatial equilibrium (Roback, 1982; Wu and Gopinath, 2008) and urban economics (Glaeser, Rosenthal, and Strange, 2010; Glaeser, Kerr, and Ponzetto, 2010) models, as well as with empirical evidence in both developed (Glaeser, Kerr, and Ponzetto, 2010) and developing countries (Naudé et al., 2008; Modrego et al., 2014). These equilibrium results could differ, however, both in the short run and also in conditions where long-run interregional migration flows might be restricted.

Chile serves as a good case for testing the implications of the model. First, the Global Entrepreneurship Monitor (GEM) Project—a large, long-term cross-country study of the drivers of entrepreneurship (Reynolds et al., 2005; Bosma and Levie, 2010)—portrays the country as one of the most entrepreneurial countries in the world, notably out-performing other countries at similar stages of development (Figure 1). While the literature emphasize the importance of a greater market size for entrepreneurship (Murphy et al., 1990; Reynolds et al., 1995; Sato et al., 2012), Figure 1 shows that Chile's entrepreneurship rate is well above that of countries with a much higher per capita GDP (United States, for example).

Second, since the mid-2000s, enhancing entrepreneurship has become a key component of Chilean economic policy, and more than 100 entrepreneurship support programs have been established in the country (Romaní et al. 2009). However, we know little about the spatial dimension of entrepreneurship in the country. Given its fundamental importance for regional economic growth and development (Acs and Armington, 2004; Audretsch and Keilbach, 2004), understanding more precisely the drivers of spatial differences in entrepreneurial activity is of critical importance for both academic and policy purposes.



Note: GDP per capita in constant U.S.\$ of 2005. Source: Own elaboration based on data from the GEM project and the World Bank, accessed through World Statistics (<http://world-statistics.org>).

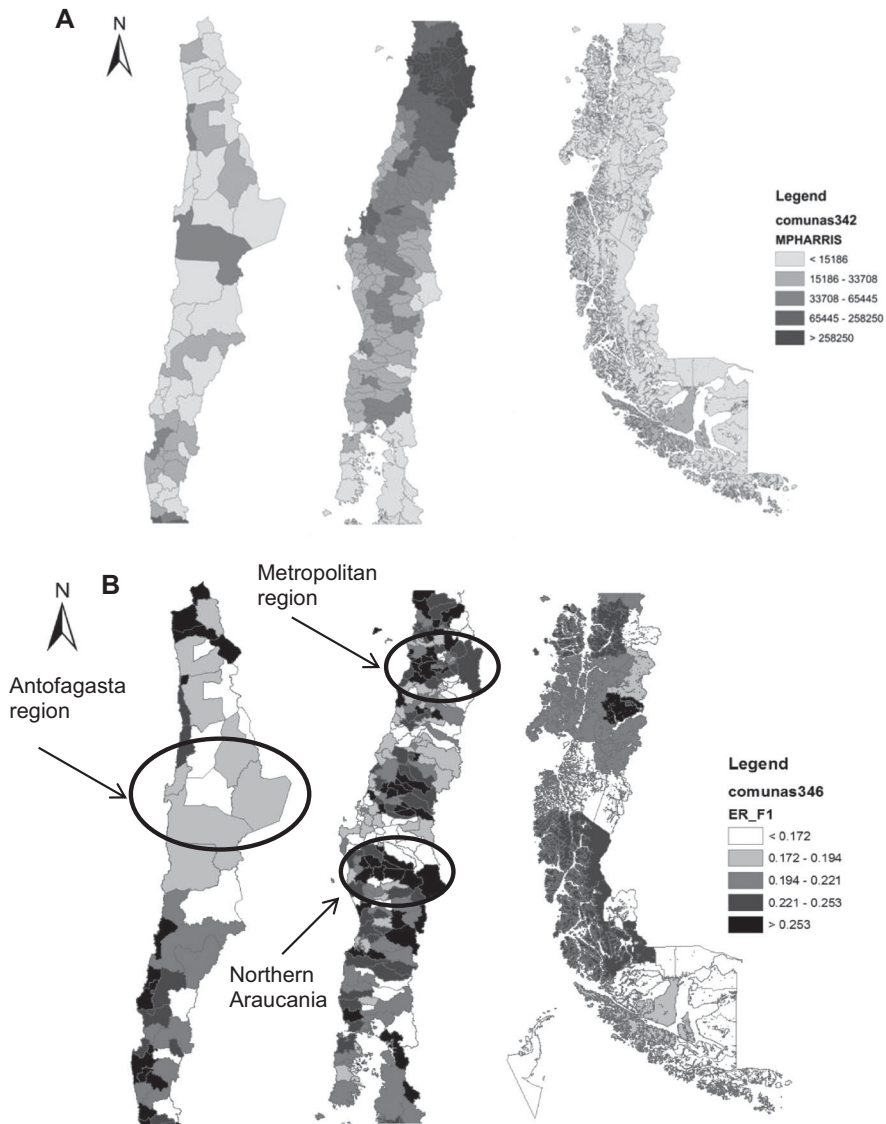
FIGURE 1: Entrepreneurial Activity and National Development (2012 or Closest).

Third, what we know is that regardless of how it is measured, entrepreneurship in Chile, as in most countries, is very unevenly distributed over space. Modrego et al. (2014) report differences in the number of firms per capita as large as 14 times between municipalities, and differences in annual business start-up rates varying from less than 1–33 new activities per 1,000 inhabitants (Modrego et al. 2015). Figures 2(a) and (b) show, as of the early 2000s, the market potential of Chilean municipalities and rates of nonfarm self-employment respectively. While there are high entrepreneurship rates in agglomerated areas, such as the Santiago Metropolitan Region, high entrepreneurship rates are also evident in many rural, remote communities, such as northern Araucanía (one of the poorest areas in the country). Apparently, the relationship between agglomeration and local entrepreneurship rates in Chile is not straightforward as well.

The following section sets out an NEG-based framework explaining the relationship between location and regional entrepreneurship rates. Section 3 presents an econometric application of the model in the previous section. Section 4 presents the data used for testing the model at the level of Chilean municipalities. Section 5 discusses the results and section 6 concludes and discusses some pending lines for future research.

2. A SPATIALLY-EXPLICIT OCCUPATIONAL CHOICE MODEL

A main feature of NEG models is providing general equilibrium solutions for the spatial distribution of economic activity, as a balance between agglomeration forces arising from increasing returns and transport costs and dispersion forces stemming from an immobile demand source (Krugman, 1991) or a nontradable sector (Helpman, 1998). Following mainstream occupational choice theories (Parker, 2009), we incorporate here another trade-off, that originating in the discrete choice each individual faces regarding becoming an entrepreneur or an employee.



Note: In both panels, Northern Chile at the left, the Center–South area in the center and the South end at the right. Source: Panel (a) Modrego et al. (2014) and panel (b) 2002 National Population Census.

FIGURE 2: Harris (1954) Market Potential and Nonfarm Self-employment Rates in Chilean *Comunas* (2002/2003).

We retain most of the assumptions in Helpman (1998). The economy comprises R regions and a mobile population of size L_r in each region r . Later we will introduce mobility and individuals' sorting across regions based on real earnings equalization (Fujita et al., 1999; Hanson, 2005).

There are two sectors in the economy: (i) a monopolistically competitive sector with increasing returns and costly trade across regions ("manufactures") and (ii) a homogeneous

good not traded across regions (H).¹ In each region, this nontradable good is supplied in fixed stock H_r , and following Helpman (1998) and Hanson (2005) we further assume that total stock of the nontradable good is equally owned by individuals in the economy, regardless of regional location. Individuals supply inelastically one unit of labor, but they choose between paid employment (l) and entrepreneurship (e). As in Lucas (1978), Evans and Jovanovic (1989), Parker (2005) and others, each individual is endowed with an intrinsic level of managerial ability (θ), drawn from a short-run fixed (continuously differentiable) regional distribution: $f^r(\theta) \rightarrow [A, +\infty]$, with $A > 0$ being the minimum level of ability in the population.

On the demand side, we introduce no innovations to the standard multi-region core-periphery model (Fujita et al., 1999). Individuals have Cobb-Douglas preferences over consumption of a composite of manufactures (m) and housing (h), constant across regions: $U = m^\mu h^{1-\mu}$, with the parameter μ indicating the budget share of manufactures. The composite variable m is a constant-elasticity-of substitution (CES) specification: $m = [\int_0^n m_i^\rho di]^{1/\rho}$, with m_i being consumption of each variety i and $\sigma \equiv 1/(1 - \rho)$ is the elasticity of substitution between varieties.

Each manufactured variety i is produced in region r at price p_{ir} and is sold across regions with trading costs. We assume iceberg transport costs, such that its price in external market s is (Hanson, 2005): $p_{is} = p_{ir}e^{\kappa d_{rs}}$, with d being the distance between markets r and s and κ a parameter measuring unit transport costs. Given assumptions, the demand for the i th differentiated variety arising from consumer’s utility maximization is expressed in terms of the Market Potential (MP) function (see Fujita et al., 1999; Hanson, 2005):

$$(1) \quad q_{ir}^d = \mu p_{ir}^{-\sigma} MP_r \equiv \mu p_{ir}^{-\sigma} \sum_{s=1}^R Y_s e^{\kappa(1-\sigma)d_{rs}} G_s^{(\sigma-1)},$$

with Y being total regional income and G_s the CES price index for region s . MP is the regional market potential, a synthetic measure of the size of surrounding markets in terms of incomes adjusted for purchasing power, negatively weighted by distance-dependent transport costs (Hanson, 2005).

On the supply side, each entrepreneur (firm) produces a different variety and the production technology entails the use of a variable input of labor, here normalized to one, plus an input of labor fixed to production levels (the source of increasing returns). Labor is paid at a regional wage, w_r , which also represents the opportunity cost of the entrepreneur, who can always earn this wage if he decides to be an employee. To introduce entrepreneurs’ heterogeneity in the production technology, we borrow from Glaeser, Kerr, and Ponzetto (2010), the idea of a managerial overhead cost $\tau > 0$, here assumed to be decreasing in the individual’s managerial ability $\theta_{ir} (\geq A)$. Accordingly, we specify the following cost function:

$$(2) \quad CT_{ir} = w_r (q_{ir} + k_{ir}) = w_r (q_{ir} + \tau \theta_{ir}^{-\gamma}),$$

with k_{ir} being the firm-specific fixed input of labor. Although these costs are fixed with respect to production levels for a given firm, the (constant across firms and regions) parameter $\gamma > 0$ represents an elasticity of fixed costs with respect to the entrepreneurs’ “managerial ability,” that is, it is a managerial efficiency parameter. Despite the extreme

¹As a reviewer correctly points out, in the short run this nontradable sector is superfluous, but it is useful to motivate differences in the cost of living in the empirical section and also we will discuss it in the context of long-run equilibrium analysis. We thank the reviewer for this observation.

simplification of the entrepreneurial function, this representation of the production technology in a market structure of potentially-unlimitedly differentiated varieties is consistent with a view of the (able) entrepreneur as the individual (more) capable of recombining resources in more efficient ways to innovate and create value (Carree and Thurik, 2003).²

Given this cost structure, the profit function of the individual entrepreneur i located in region r is:

$$(3) \quad \pi_{ir} = (p_{ir} - w_r) q_{ir} - w_r \tau \theta_{ir}^{-\gamma}.$$

The monopolistic entrepreneur in the differentiated sector sets prices such that marginal benefit equals marginal cost; here, this yields the constant mark-up rule. Given marginal costs independence of individual ability, the price is the same for each variety produced in the region r as in the standard NEG core-periphery model:

$$(4) \quad p_r = \frac{\sigma w_r}{\sigma - 1},$$

with σ also being the price elasticity of the demand for any variety produced in region r , with $\sigma > 1$ due to the maximizing monopolist's operating in the elastic part of the demand curve.

Embedding the pricing rule (4) and the demand equation (1) into the profit function (3) this latter can be rewritten as:

$$(5) \quad \pi_{ir} = \phi MP_r w_r^{1-\sigma} - w_r \tau \theta_{ir}^{-\gamma},$$

with $\phi = \mu \sigma^{-\sigma} (\sigma - 1)^{\sigma-1}$.

Characterizing the Equilibrium in the Short Run

We wish to examine equilibrium conditions when factors within a region can move freely; in particular, we define the short-run equilibrium as wages and profits that make worker indifferent to becoming entrepreneurs. Later we will address the long-run equilibrium, which we define as labor flows across regions based on differential real earnings (e.g., Fujita et al., 1999; Hanson, 2005).

Most previous NEG-based models incorporating entrepreneurship rely on the zero-profits entry condition to derive the equilibrium relationships. However, as pointed by Glaeser, Rosenthal, and Strange (2010, p. 4), the zero-profits equilibrium “*takes entrepreneurship out of the model*,” by implicitly assuming a horizontal supply of entrepreneurs across space. To incorporate the individual decision of starting a business in this spatial economic model, we follow an occupational choice setting. As in Lucas (1978), Evans and Jovanovic (1989), Parker (2005) and others, workers choose entrepreneurship as long as the returns to self-employment are at least as high as those of hired labor. The short-run (long-run) free-entry occupational choice equilibrium condition for all ability levels choosing entrepreneurship is then:

$$(6) \quad \pi_{ir} \geq w_r \quad \forall r.$$

²One could think of alternative ways of incorporating differences in managerial ability. For instance, occupational choice models usually assume that greater ability translates into greater productivity via shifts outward in the production function (e.g., Lucas, 1978; Evans and Jovanovic, 1989). Alternatively, one could assume that managerial ability also entails marginal cost-efficiency. Both approaches add algebraic complications, but are inessential to the main thrust of the model, which is that endogenous entrepreneurs' opportunity costs (wages) offset pro-entrepreneurship market-size effects.

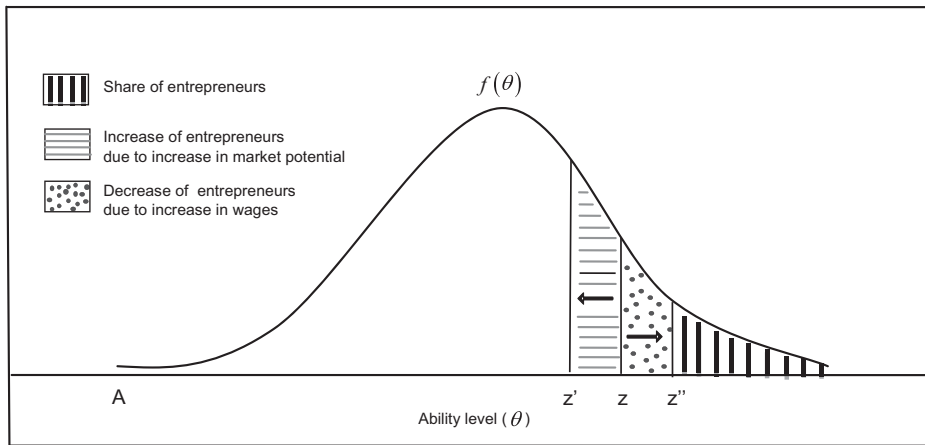


FIGURE 3: Partial Effects of Increases in Market Potential and Domestic Wages on the Short-Run Regional Equilibrium Level of Managerial Ability.

At the ability level where entrepreneurial income is equal to wages produces a *regional marginal ability function* (z_r) defining the point at which the marginal entrepreneur located in region r breaks even, that is, the point at which incomes from the entrepreneurial undertaking equals the total costs (including production and the opportunity cost of not being an employee). This function is decreasing in the regional MP and increasing in domestic wages. By inserting equation (5) into (6) at the equilibrium level of ability yields

$$(7) \quad z_r = \tau^{1/\gamma} (\phi MP_r w_r^{-\sigma} - 1)^{-1/\gamma}.$$

Equation (7) states that, *everything else constant*, in areas of higher market potential, a lower level of managerial ability is needed to succeed in business, since interregional demand linkages provide opportunities for marginal (less able) entrepreneurs to make profits. The effect of an increase in the market potential is illustrated in Figure 3 as the shift to the left in the cut-off level of ability (from z to z'). The model thus retains the idea of larger demand as a favorable condition for entrepreneurship (Sato et al., 2012; Modrego et al., 2014). But at the same time, it casts a positive *partial effect* of domestic wages over z . Since the individual can access this wage (which is also what the entrepreneur must pay his employees) by choosing being an employee, when wages are higher, less able potential entrepreneurs will rationally opt to be paid employees (Lucas, 1978; Parker, 2005). This can be seen in Figure 3 as the displacement of the cut-off level of ability to the right, from point z to z'' . Given that z_r is nonnegative, then in equilibrium wages would be such that: $(\phi MP_r)^{1/\sigma} \geq w_r$, and so assuring an interior solution with both entrepreneurs and workers in the region.³

In this model, relationship (7) establishes a subtle but crucial difference from Sato et al. (2012), where regional variations in the opportunity cost of entrepreneurship (wages) are taken as exogenously determined. In this present treatment, we instead make the entrepreneur’s opportunity cost endogenous, following one of the main implications of the NEG core–periphery model, which is the positive dependence of regional wages on regional market potential (Fujita et al., 1999; Hanson, 2005).

³We thank an anonymous reviewer for motivating this characterization of the equilibrium solution.

Our strategy is to develop the labor market equilibrium to check whether this NEG's core relationship is affected by the introduction of occupational choice elements in the model. We also assess the net effect of a shock in the MP on the regional supply of entrepreneurs. These net effects are both direct through business profits and indirect through equilibrium wages.

The short-run regional supply of entrepreneurs is the number of people with a level of ability above the cut-off point:

$$(8) \quad e_r = L_r \int_{z_r}^{+\infty} f^r(\theta) d\theta = L_r [1 - F^r(z_r)].$$

Analogously the regional short-run supply of employees (l^s_r) is

$$(9) \quad l^s_r = L_r F^r(z_r).$$

Given assumed technology, the firm-level demand for labor (l^d_{ir}) is

$$l^d_{ir} = q_{ir} + \tau \theta_{ir}^{-\gamma} = \omega MP_r w_r^{-\sigma} + \tau \theta_{ir}^{-\gamma},$$

with $\omega = \phi/(\sigma - 1)$.⁴

This means, given the way the managerial ability enters the cost function, that less able entrepreneurs in the region will demand more labor per output, generating higher unit costs. In the absence of agglomeration externalities on production (as in Sato et al., 2012), this means that *keeping everything else constant*, regions with higher market potential will have more firms, of larger average size and lower average labor productivity.

In turn, the regional demand for labor is given by (see Lucas, 1978):

$$(10) \quad l^d_r = L_r \int_{z_r}^{+\infty} l^d_{ir}(\theta) dF^r = L_r (1 - F^r(z_r)) \left[\omega MP_r w_r^{-\sigma} + \bar{k}^r(z_r) \right].$$

This is equivalent to the total number of entrepreneurs (firms) in equation (8) (the term outside brackets), multiplied by the average labor input per firm (the term in brackets). The latter is comprised of the (constant across firms in the region) variable labor input (first term in brackets) and the average fixed labor input per firm in the region ($\bar{k}^r(z_r)$). This regional labor demand is increasing in the MP and decreasing in domestic wages (as are both the demand faced by each firm and the number of firms).

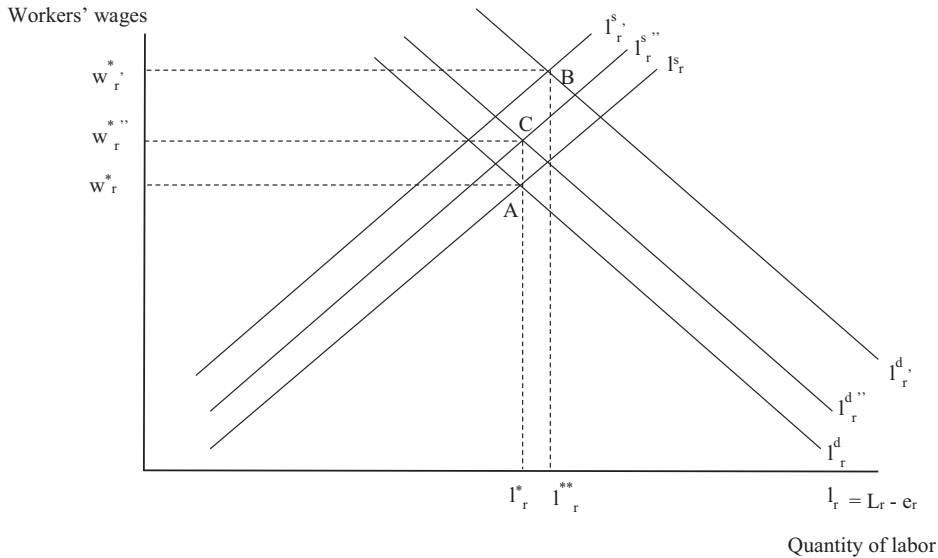
Equalization of the regional labor supply (9) and the regional demand for workers (10) defines the equilibrium regional wages as a nonlinear function of the MP:

$$(11) \quad w_r = \left[\left(\frac{F^r(z_r)}{(1 - F^r(z_r))} - \bar{k}^r(z_r) \right) \frac{1}{\omega MP_r} \right]^{-1/\sigma}.$$

An implication of equation (11) is that *ceteris paribus*, a higher share of entrepreneurs translates into higher wages, because more entrepreneurship increases labor demand while reducing salaried-labor supply. Equations (7) and (11) solve the model for the regional equilibrium number of entrepreneurs (and employees) and regional wages as a function of the MP. Two main results, which are also testable hypotheses of the model, that arise from the short-run equilibrium conditions are:

Hypothesis 1: Regional wages are increasing in the market potential. More specifically, as in the standard NEG core-periphery setting, the elasticity of wages with respect to changes in the MP is the inverse of the elasticity of substitution.

⁴This result is an application of Hotelling's lemma to the profit function as given in equation (5).



Note: Greater market potential shifts labor demand rightward, from l_r^s to $l_r^{s'}$, raising wages from w_r^* to $w_r^{*'}$ (from point A to point B). Increased wages draws marginal entrepreneurs to switch back into paid employment, thus increasing the supply of labor and decreasing demand (from point B to point C). The net effect of these changes is a new equilibrium with the same share of entrepreneurs and paid workers but with higher wages, as point C.

FIGURE 4: The Short-Run Regional Labor Market Equilibrium—Effects of an Increase in the Market Potential.

Hypothesis 2: In the short run, before wage differentials trigger immigration, the equilibrium share of entrepreneurs is invariant to shocks in the MP, as the higher profit-effect is completely offset by the opportunity-cost effect.

We leave to appendix 1 and 2 the formal proof of hypotheses 1 and 2, as an exercise in comparative statics motivated by a marginal change in the MP. Still, one simple way of analytically verifying the second hypothesis is to apply equation (11) defining the local labor market equilibrium wage to equation (7) defining the regional marginal ability function to note that the indifference level of ability is determined by the regional distribution of abilities and not reference to market potential or other variables:⁵

$$z_r = \tau^{1/\gamma} \left[(\sigma - 1) \left(\frac{F^r(z_r)}{1 - F^r(z_r)} - \bar{k}^r(z_r) \right) - 1 \right]^{-1/\gamma}.$$

To develop the intuition of these results here, Figure 4 depicts the short-run (before wage-driven immigration) adjustments in the labor market triggered by a positive shock in regional MP, such as the building of a new road or a new (spatial targeted or not) public transfer program. The initial effect of an increased MP is a decrease in the regional break-even level of ability. This provokes a shift to the left of the supply of workers, as some individuals previously below the cut-off level of ability switch into entrepreneurship. Incumbent entrepreneurs now individually demanding more labor and new entrepreneurs demanding increasingly more labor shift the demand curve to the right, forcing wages to go

⁵We would like to thank a *JRS* reviewer for suggesting this interpretation.

up to restore the equilibrium (from point A to point B). Wages will increase to the original break-even level of ability, making marginal entrepreneurs switch into paid employment. This shifts the supply for labor to the right and the demand to the left. As demonstrated in the appendix 1, the net effect of these changes is a new equilibrium with the same share of entrepreneurs and employees but with higher wages, as point C in the figure. In fact, the second hypothesis also holds in the long run, as will be demonstrated below.

In this NEG-based model with only one production factor, the increase in wages is the only way the surplus generated by the increase in MP is shared between entrepreneurs and employees. Additional rent is bid into both the entrepreneurial profits and into the wages of workers. The relative gains of the two groups depend on the size of the groups, but as the appendix 2 demonstrates, all participants regardless of ability (θ) gain proportionally the same amount ($1/\sigma$). The greater is the elasticity of demand, the smaller would be the potential mark-up of prices over marginal costs and smaller would be the total rents available to be shared between both entrepreneurs and employees.

*Long-Run Equilibrium Results When Labor Is Mobile Across Regions*⁶

In the long run, both entrepreneurs and employees can migrate across regions. To characterize the long-run equilibrium, we examine the three short-run endogenous variables of interest in the previous section: entrepreneur profits at every ability level (equation 5), the regional marginal ability function (equation (7)), and regional nominal wages, which equilibrate the local labor market (equation (11)). These are functions of a long-run endogenous variable: the Market Potential function (defined in equation (1)). Market potential is, in turned, defined by the regional price index (see Fujita et al., 1999; Hanson, 2005):

$$(12) \quad G_r = \left[\sum_{s=1}^R n_s \left(\frac{\sigma w_s}{\sigma - 1} e^{k d_{rs}} \right)^{1-\sigma} \right]^{\frac{1}{1-\sigma}},$$

with n_s being the number of varieties of manufactures, which is equal to the number of firms, and to the number of entrepreneurs in regional market s :

$$n_s = e_s = L_s \int_{z_s}^{\infty} f^s(\theta) d\theta.$$

An additional (short-run and long-run) equilibrium condition is the clearing of the nontradable market with inelastic supply (total payments equating total expenditures) (Hanson, 2005):

$$(13) \quad p_r^h H_r = (1 - \mu) Y_r, \quad \forall r,$$

such that total regional incomes are defined as

$$(14) \quad Y_r = L_r \left(F^r(z_r) w_r + \int_{z_r}^{+\infty} f^r(\theta) \left(\phi MP_r w_r^{1-\sigma} - w_r \tau \theta r^{-\gamma} \right) d\theta + \frac{1}{L} \sum_{s=1}^R p_s^h H_s \right).$$

The first term in the parenthesis of equation (14) (when multiplied by L_r) represents the wages times the number of employees; the second is the entrepreneurs' mean earnings

⁶This subsection was absent in a previous version of the paper and it was motivated by the comments of three *JRS* reviewers to whom we are extremely grateful.

times the number of entrepreneurs, and the third is the income from the nontradable sector.

Some long-run equilibrium results fall out of assuming an interregional labor market equilibrium with costless migration, which means that no individual, regardless of the ability level, would have an incentive to move between regions and/or occupations. Following Hanson’s (2005) treatment of labor market equilibria with costless migration, real wages depend on nominal wages deflated by a regional price index (defined by the consumers’ expenditure function) and equal across regions:⁷

$$(15) \quad \frac{w_r}{(p_r^h)^{1-\mu} G_r^\mu} = \frac{w_s}{(p_s^h)^{1-\mu} G_s^\mu}, \quad \forall r, s.$$

Similarly, in this present model, the real profits of entrepreneurs (individuals at or above the cut-off level of ability) must also be equal across regions at every ability level, otherwise entrepreneurs would migrate, either to be entrepreneurs in another region or to be an employee:

$$(16) \quad \frac{\pi_{ir}}{(p_r^h)^{1-\mu} G_r^\mu} = \frac{\pi_{is}}{(p_s^h)^{1-\mu} G_s^\mu} \quad \forall \theta_i \geq z_t \text{ and } \forall t = r, s.$$

Note that free-entry occupational choice equilibrium (equation (6)) means also that real earnings are equal across regions and occupations at the indifference ability level of the marginal entrepreneur:

$$(17) \quad \frac{\pi_r^z}{(p_r^h)^{1-\mu} G_r^\mu} = \frac{w_r}{(p_r^h)^{1-\mu} G_r^\mu} = \frac{w_s}{(p_s^h)^{1-\mu} G_s^\mu} = \frac{\pi_s^z}{(p_s^h)^{1-\mu} G_s^\mu}, \quad \forall r, s.$$

Noting that profits at all ability levels are proportional to wages:

$$\pi_{ir} = \phi MP_r w_r^{1-\sigma} - w_r \tau \theta_i^{-\gamma} = w_r \cdot (\phi MP_r w_r^{-\sigma} - \tau \theta_i^{-\gamma}),$$

condition (16) combined with the symmetry of production technology across regions entails that, in the long run, all entrepreneurs of the same ability level θ earn the same premium above wages:

$$\phi MP_r w_r^{-\sigma} - \tau \theta^{-\gamma} = \phi MP_s w_s^{-\sigma} - \tau \theta^{-\gamma} \Rightarrow MP_r w_r^{-\sigma} = MP_s w_s^{-\sigma}.$$

And in particular, the equilibrium ability level is the same for all regions:

$$(18) \quad z^r = z^s = \bar{z}, \quad \forall r, s.$$

This latter result suggests another observable consequence of the theory, at least in the long run: in equilibrium, the relative *nominal* wages between two regions are proportional to the ratio of regional market potentials, as also is the cost of living ratio:

$$(19) \quad \frac{w_r}{w_s} = \left(\frac{MP_r}{MP_s} \right)^{1/\sigma} = \left(\frac{p_r^h}{p_s^h} \right)^{1-\mu} \left(\frac{G_r}{G_s} \right)^\mu.$$

⁷As remarked by one *JRS* reviewer, recent NEG models (e.g., Südekum, 2008) emphasizing the role of amenities, propose a spatial equilibrium sorting based on the equalization of indirect utilities. Because our focus is placed on the role of NEG’s market potential function in conditioning the regional supply of entrepreneurs, we resort to the standard real wage equalization mechanism followed by Fujita et al. (1999) and Hanson (2005), among others. In absence of amenities, the principle of deflating nominal wages is exactly the same whether one equates real wages or indirect utilities. We thank the reviewer for bringing up this point.

Relationship (19) reflects the standard type of interregional, economic-geography equilibrium model with the result that firms must compensate workers for living in a costly (or unpleasant/congested) place (Roback, 1982; Wu and Gopinath, 2008). Firms do so by offering higher wages, which is feasible since firms enjoy some location advantage, in this case greater profits due to market proximity (a greater market potential).

Finally, given that the marginal abilities are equal across regions one notes that in this model, long-run differences in entrepreneurship rates are, given the distributions of abilities in regions, independent of all other variables:

$$(20) \quad \frac{F^r(\bar{z})}{1 - F^r(\bar{z})} - \frac{F^s(\bar{z})}{1 - F^s(\bar{z})} = \bar{k}^r(\bar{z}) - \bar{k}^s(\bar{z}) = \int_{\bar{z}}^{\infty} \tau \theta^{-\gamma} f^r(\theta) d\theta - \int_{\bar{z}}^{\infty} \tau \theta^{-\gamma} f^s(\theta) d\theta.$$

This result can be verified by inserting equation (11) into (7) and using result (17). Long-run equilibrium differences in entrepreneurship rates depend in differences in average fixed costs between regions. In other words, regions with higher average fixed costs of doing business would have lower entrepreneurship rates, a result that is also present in urban economic models of entrepreneurship (Glaeser, Kerr, and Ponzetto, 2010; Glaeser, Rosenthal, and Strange, 2010). For instance, if the regional density distributions of entrepreneurial ability are equal, then average fixed costs and entrepreneurship rates will also be equal. This is not say that the distributions of abilities by region are themselves not functions of the more basic NEG parameters, such as the stock of the nontradeable good, H_r , or the transport costs between regions, κd_{rs} . On the contrary, such distributions depend on the nature and patterns of migration behavior in the context of the geographical features relating to the supply of housing and infrastructure, all of which themselves shape market potential. Based on this model, there are no *a priori* conclusions as to what these patterns might look like in the long run.

As noted above, a positive shock to MP in a region delivers proportional gains across all ability levels, which in the long run would induce immigration from other (less fortunate) regions. If one were to assume that migration costs are independent of the entrepreneurial ability level, a reasonable assumption in the light of Lazear's "jack-of-all-trades" hypothesis (Lazear, 2004, 2005), this would imply immigration from all talent levels.⁸ In effect, migration from other regions would be a *pure scale effect*, as if peeling a layer from the talent distribution of the rest of the world and laying it on top of that of the region enjoying the increase in market potential.⁹ If, on the contrary, one expects a negative correlation between entrepreneurial ability and the cost of migration, the result would be agglomeration with increased entrepreneurship rates (a dynamic that is inconsistent with stylized facts indicating a negative correlation between urban density and self-employment rates, see Behrens and Robert-Nicoud, 2015).

The exactly compensating profit and opportunity cost effects driven by changes in market potential may seem, at first glance, an intuitively puzzling result, given the

⁸Although it is usually assumed that entrepreneurs are more educated people, recent profiles show that self-employed and employees in Chile are not very different in terms of formal schooling (10.1 and 11.5 years respectively) (Modrego, Paredes, and Romani, 2015).

⁹Given an even stronger assumption of similar initial ability distributions across regions, and the similar proportional gains across all talent levels that would induce immigration from elsewhere, the final equilibrium distribution of ability in the beneficiary region would not change. The absolute number of entrepreneurs and employees would increase, driving down local entrepreneurial real rents and wages; and the absolute number of entrepreneurs and employees elsewhere would decline, driving up real rents and wages. A final equilibrium, long-run result would be one region with a larger population but with the same shares of workers and entrepreneurs, and a sharing across all regions of the gains due to the region's enhanced market potential.

theories and empirical evidence supporting the positive relationship between market size or market potential and entrepreneurship (Audretsch and Fritsch, 1994; Sato et al., 2012) and business density (Modrego et al., 2014). However, as remarked by Behrens and Robert-Nicoud (2015), whether the share of entrepreneurs increases or decreases with market size or density strongly depends on the modelling choices and empirical evidence is ultimately inconclusive. In this present application, the result directly stems from the model’s structure: the constant mark-up rule, the production technology with only fixed costs depending on the managerial ability, and the free-entry equilibrium condition of entrepreneurs given by the equalization of business profits and wages. Combined, these conditions yield a proportional relationship between the opposing incentives conditioning the selection into entrepreneurship, which in turn leads to these symmetric shifts in the labor market and to long-run differences independent of regions’ MP.¹⁰ These predictions can be interpreted as consistent with some empirical evidence from developed economies discarding the hypothesis of places with abnormal business returns as a main explanation for the spatial variation of entrepreneurship (Glaeser, Kerr, and Ponzetto, 2010). As shown in the next section, the compensating rents and wage effects fit well with the case of municipal entrepreneurship rates in Chile.

3. AN EMPIRICAL APPLICATION

One important implication of the model developed above is the offsetting effect of location in relation to market with regard to determining the short-run equilibrium supply of entrepreneurs in a region. In this section, we present an econometric strategy that uses short-term differences in municipal entrepreneurship rates in Chile to test the short-run equilibrium hypotheses presented in the previous section.

We start from the regional supply of entrepreneurs (equation (8)). To arrive at a specific functional form, we follow Lucas (1978) and assume a Pareto distribution function for the entrepreneurial ability: $F(\theta) = 1 - A^n\theta^{-\eta}$, with $\eta > 0$ being a parameter that here measures the elasticity of the supply of entrepreneurs to the managerial ability.

Expressing the equilibrium supply of entrepreneurs (8) in rates and totally differentiating yields

$$(21) \quad d \ln \left(\frac{e}{L} \right)_r = -\eta \cdot d \ln z_r.$$

Taking the total differential of the regional marginal ability function (equation (A.1) in the appendix 1) and rearranging terms:

$$(22) \quad d \ln(z)_r = (d \ln(MP)_r - \sigma d \ln(w)_r) \left(-\frac{\phi MP_r w_r^{-\sigma}}{\gamma (\phi MP_r w_r^{-\sigma} - 1)} \right) \\ \approx -\frac{1}{\gamma} d \ln(MP)_r + \frac{\sigma}{\gamma} d \ln(w)_r,$$

¹⁰There are ways to modify this sharp result of the model. One is to assume an additive utility function that includes a subjective element of personal satisfaction of being self-employed, as in Blanchflower and Oswald (1998). In such case, the individual equates the sum of pecuniary and nonpecuniary utility of entrepreneurship with the wage income utility of being an employee. In equilibrium, marginal entrepreneurs will end up earning less than the regional wage and the opportunity-cost effect will more than compensate the entrepreneurial-utility effect, leading to lower entrepreneurship rates in presence of an increase in the MP. While appealing, such an approach would create an identification problem in the empirical strategy we devise to test the trade-offs driven by location over entrepreneurship, because it introduces an additional unobservable parameter. Another option is to include a spatially-variant start-up cost, which would lead to the opposite equilibrium result. We lack of an appropriate proxy in the empirical analysis.

which is a good approximation provided MP_r is sufficiently large relative to w_r .¹¹

Combining equations (21) and (22):

$$(23) \quad d \ln \left(\frac{e}{L} \right)_r \approx \frac{\eta}{\gamma} d \ln (MP)_r - \frac{\eta\sigma}{\gamma} d \ln (w)_r.$$

Equation (23) reflects the equilibrium relationship as derived from the theoretical model. However, in practice observed entrepreneurship rates likely deviate from such a state, so we propose the following dynamic linear reduced form:

$$(24) \quad \ln \left(\frac{e}{L} \right)_{rt} = \beta_0 + \beta_1 \ln \left(\frac{e}{L} \right)_{rt-1} + \beta_2 \ln (MP)_{rt} + \beta_3 \ln (w)_{rt} + \lambda \ln (\mathbf{x})_{rt} + \nu_r + \delta d_t + \varepsilon_{rt}.$$

In equation (24), β_1 captures path dependencies in regional entrepreneurship dynamics (Fritsch and Mueller, 2007; de Blasio and Nuzzo, 2010; Andersson and Koster, 2011). ν_r accounts for time-invariant features of regional economies potentially affecting the supply of entrepreneurs, such as institutions (Saxenian, 1994; Boettke and Coyne, 2003) or regional entrepreneurial culture (Davidson and Wiklund, 1997). In terms of the model developed above, these fixed effects are accounting for region-specific parameters shaping the different regional ability distribution functions ($f^r(\theta)$). The vector \mathbf{x} contains other economic and demographic variables suggested in the literature as potentially conditioning the supply of entrepreneurs (Reynolds et al., 1995; Blanchflower and Oswald, 1998), and are discussed in the data section. Term d_t are time dummies capturing transitory shocks common to all regions.

Time-differencing (24) eliminates the regional fixed effect, such that the final estimation equation is

$$(25) \quad \Delta \ln \left(\frac{e}{L} \right)_{rt} = \beta_1 \Delta \ln \left(\frac{e}{L} \right)_{rt-1} + \beta_2 \Delta \ln (MP)_{rt} + \Delta \beta_3 \ln (w)_{rt} + \lambda \Delta \ln (\mathbf{x})_{rt} + \delta d_t + \Delta \varepsilon_{rt}.$$

Equation (25) is the empirical counterpart of our comparative statics exercise in Figure 4 and appendix 1. It models the process of net entry into entrepreneurship as a response to changes in the regional market potential and regional wages.

Our main tests of hypothesis are the following:

$$(26a) \quad \beta_2 > 0, \quad \beta_3 < 0.$$

A more stringent implication arising from relationship (23) is

$$(26b) \quad -\frac{\beta_3}{\beta_2} = \sigma > 1,$$

consistent with the monopolist’s maximizing behavior. Previous estimates of parameter σ for Chile being around 1.5–1.8 (Modrego et al., 2014) provide a benchmark for further validation of the model.

To compute regional market potential, we follow Hanson (2005) and make use of the equilibrium sorting of workers, given by equalization of real earnings across regions (equation (16)). Solving for G_s in (15) and raising the resulting expression to $(\sigma - 1)$, yields (see Hanson, 2005):

$$(27) \quad G_s^{\sigma-1} = C w_s^{\frac{\sigma-1}{\mu}} \left(P_s^h \right)^{\frac{(\sigma-1)(\mu-1)}{\mu}}.$$

For this empirical application, we follow the convention of conceptualizing the non-tradable sector (H) as housing services, such that p_s^h indicates housing prices. As discussed

¹¹In our dataset $\frac{\phi MP_r w_r^{-\sigma}}{\phi MP_r w_r^{-\sigma} - 1}$ ranges from 1 to 1.008.

in Hanson (2005), C is a function of constant terms across regions and it equals the left-hand side of Equation (15). Replacing (27) in (1) and omitting the constant C yields the following market potential function used in the empirical application:

$$(28) \quad MP_r = \sum_{s=1}^R Y_s e^{\kappa(1-\sigma)d_{rs}} w_s^{\frac{\sigma-1}{\mu}} \left(p_s^h \right)^{\frac{(\sigma-1)(\mu-1)}{\mu}}.$$

Note that our main theoretical restriction entails the estimation of the elasticity-of-substitution parameter, σ , which is at the same time part of the market potential function (28), one of our two principal covariates. Instead of relying on nonlinear techniques that may carry some problems in the context of instrumental variables estimation (see Mion, 2004), we follow Fingleton (2006), Fingleton and Fisher (2010) and others in taking previous estimates of structural NEG parameters when computing the MP function in each region. We choose the following values estimated for Chile by Modrego et al. (2014) (average of columns 1–4 in Table 2): $\mu = 0.87$, $\kappa = 40$ and $\sigma = 1.6$. If the proposed theory fits the Chilean data, one should expect the term $-\frac{\beta_3}{\beta_2}$ to not be different (statistically speaking) to this value of σ assumed for MP's calculation. Yet, since this *a priori* value chosen for σ may be considered small given other estimates reported in the literature (Mion, 2004; Hanson, 2005; Fingleton, 2006; Fingleton et al., 2010), we perform a robustness check by recomputing the MP variable for different elasticities of substitution, and then checking if relationship (26b) is still verified with these new MP's.

Given the endogeneity of our two main regressors (MP and wages) and due to the absence of good external instruments, model (25) is estimated by means of the generalized method of moments (GMM) estimator proposed by Arellano and Bond (1991) (A-B), implemented in Stata by Roodman (2009). This is currently a standard econometric technique for estimating dynamic panel models, and a powerful tool for reduced-form NEG equations with endogenous regressors (Mion, 2004). Despite that, there are still very few applications of the A-B method in empirical NEG studies. The estimates were carried using a two-step GMM estimator, which is robust in presence of heteroscedasticity and the Windmeijer (2005) finite sample correction was applied in order to avoid the well-known small-sample biases of coefficient's standard errors (Arellano and Bond, 1991).

4. THE DATA

Our data source are ten rounds of the Chilean National Survey of Socioeconomic Characterization (CASEN) conducted in the years: 1990, 1992, 1994, 1996, 1998, 2000, 2003, 2006, 2009, and 2011.¹² The variables were collected at the level of *comunas* (municipalities), the lowest administrative unit in the Chilean administrative hierarchy. There were 342 *comunas* in 1990.¹³ Of these 342 *comunas*, the variables were built for a group for which CASEN's design yields a representative sample in each round (called "self-represented *comunas*"). For such *comunas*, CASEN provides *comuna*-level weights for expanding the sample to the *comuna*'s total population (see Mideplan, 2005), and such weights were used for expanding all variables built here. As such set of *comunas* varies in each round of the survey, we ended up with an unbalanced panel of 2,305 observations, which include observations for 331 of the 342 *comunas* in the country.

Our dependent variable (entrepreneurship rates in the theoretical model) is the proportion of self-employed plus employers to the working population, consistent with our

¹²There was also a round in 1987, which is currently unavailable in CASEN's website: http://observatorio.ministeriodesarrollosocial.gob.cl/casen_obj.php

¹³New *comunas* created during the period of analysis were treated as being part of the original *comuna*.

occupational choice framework and also a traditional metric in empirical analysis of entrepreneurial activity (see Parker, 2009). It excludes self-employed and employers in the natural resources (NNRR) primary sector (farmers, artisan fishermen), as entrepreneurship is usually associated with nonfarm proprietorship (e.g., Goetz and Rupasingha, 2009). Employees in the NNRR primary sector were also excluded as a model with a tradable industry under increasing returns and monopolistic competition is largely irrelevant in this context (Fingleton, 2006).

Following the standard practice, we calculate wages as the average income of the main occupation of employees in the comuna (Paredes, 2013). To avoid obvious aggregation biases, for all self-represented comunas in each CASEN round we calculate the municipal market potential in relation to a set of 51 comunas (the set s in equation 20) that are self-represented in all rounds in the 1990–2011 period. Although this way of constructing the MP variable comes at the cost of some measurement error, we believe that our empirical metric is still a good proxy of the true market potential. This set of 51 comunas encompasses most of the major local markets, including all the regional capital cities and most of the largest provincial capitals. Regarding the MP variables, regional incomes were taken as the sum of total (from all sources) incomes of all households in the comuna and housing prices were proxied by the average imputed housing rental value in the comuna. Monetary variables were expressed in monthly thousands Chilean pesos (base December 2009), deflated by the consumer price index reported by the Central Bank of Chile.¹⁴ A matrix of distances between municipal capital cities was taken from the Ministry of Public Works and distances were expressed in thousands of kilometers.¹⁵

To test for the robustness of results in the base-specification, we include sociodemographic and economic variables shown as conditioning the regional supply of entrepreneurs. We first consider the share of municipal population above 24 years old with complete university (tertiary) studies, with the *a priori* expectation being an entrepreneurs' supply increasing, *ceteris paribus*, with greater levels of education, as more skilled individuals are more likely to succeed in self-employment (Robinson and Sexton, 1994; Parker, 2005). We also include the percentage of people in mid-labor career in the *comuna* (i.e., between 35 and 45 years old), a variable indicating the share of people at the moment of the life cycle supposedly most favorable for starting the own business. Results for the U.S. indicate that this variable is positively related to local business start-up rates (Reynolds et al., 1995). We also include a Herfindhal index of economic diversification, based on workers occupations according to the one-digit International Standard Classification of Occupations (ISCO). It ranges from zero (total specialization) to 100 (perfect diversification) and it is intended at capturing productivity and innovation-enhancing diversification (also called *Jacobs*) externalities (see Glaeser et al., 1992). Finally, the proportion of women in the population, since women have been regarded as economically and institutionally limited to self-employment (Hundley, 2001).

5. RESULTS

Exploratory Analysis

Figure 5 depicts the three main relationships derived from the theoretical model. Panel (a) shows a weak but positive association between changes in the market potential

¹⁴<http://si3.bcentral.cl/Siete/secure/cuadros/home.aspx> Last accessed: March 4, 2014.

¹⁵<http://www.vialidad.cl/productosyservicios/paginas/distancias.aspx> Last accessed: December 19, 2013.

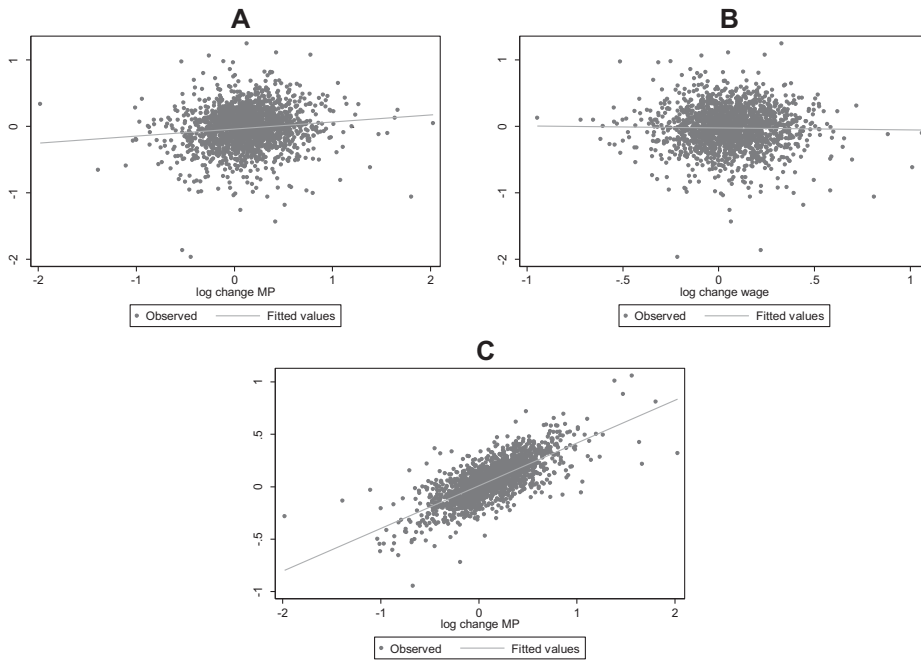


FIGURE 5: Changes in Market Potential, Wages, and Entrepreneurship Rates in Chilean *Comunas* (Logs).

and changes in self-employment rates in Chilean comunas ($\rho = 0.119$). This simple correlation is small, although not inconsistent, when compared to the predictions of previous NEG models of entrepreneurship (Sato et al., 2012; Modrego et al., 2014) and also when compared to empirical results from several countries pointing at a robust positive correlation between agglomeration and entrepreneurial activity (Audretsch and Fritsch, 1994; Sato et al., 2012).

Panel (b) in turn, indicates a virtually nonexistent simple correlation between wage variations and changes in self-employment rates ($\rho = -0.018$), which goes against the mechanism underlying most occupational choice models and vast empirical evidence from different countries (Parker, 1996; Belso Martínez, 2005). On the other hand, Panel (c) shows a strong positive association between changes in market potential and changes in wages ($\rho = 0.736$). This result is consistent with hypothesis 1 above (that regional wages are increasing in market potential) and with empirical evidence in countries as diverse as the U.S. (Hanson, 2005), Italy (Mion, 2004), and China (Hering and Poncet, 2010).

Taken as a whole, these three pieces of information suggest that the simple correlations between regional self-employment rates, market potential and wages may be obscured by their mutual interdependencies. Market potential actually seems to exert part of its influence over regional self-employment rates indirectly through wages, what possibly explains the nil (instead of strongly negative) relationship between changes in wages and entrepreneurship rates in Figure 5(b). Further, if the (negative) opportunity cost-effect driven by higher wages offsets the (positive) profit-effect driven by spatial demand linkages, as the model above emphasizes, one would expect patterns similar to those in Figure 5.

These underlying trade-offs can be better assessed through estimation of regression parameters reflecting the *ceteris-paribus* partial effect of each covariate on the dependent variable. Along with parameter restrictions (26a) and (26b), this is a demanding statistical test of the proposed theory. We now present estimates of our econometric specification.

Dynamic Panel Estimates

As noted above, the observations are irregularly spaced (with two- or three-year intervals, depending on the period). This varying interval length introduces several problems in the context of dynamic panel models (see Millimet and McDonough, 2013). Millimet and McDonough (2013) evaluate a range of estimators for panels with irregular spacing, including the A-B estimator, and conclude that they are all inconsistent in presence of endogenous regressors (see table 3 in Millimet and McDonough, 2013). Wages and the market potential function are jointly determined in the theoretical model proposed here, so their exogeneity in the estimated model would be unwarranted. For these reasons, we estimate the dynamic panel equation (25) for two regularly spaced periods: 1990–2000 (two years between rounds) and 2000–2009 (three-year intervals). Nevertheless, we also provide estimates for the whole 1990–2011 period with irregular intervals, while acknowledging the inconsistency of the A-B estimator in this case, to assess the implications for our results of not considering the unequal spacing between time spans.

Following the standard practice in applied studies using dynamic panel estimation, we use second-order lags in levels and differences to instrument the endogenous variables. In particular, we use lagged and twice-lagged *differences* of the market potential index and wages, and the second-order lags of the dependent variable in *levels* (which means, given the time spacing of the panel, entrepreneurship rates of the previous four or six years depending on the estimation period). One notes that the market potential measure varies within each group by considerably less than the measure's within-group averages vary across the panel; and so one might suspect some persistence to the between-group endogeneity of the lagged levels of the variable as instruments. However, the underlying conceptual argument for persistence of the market potential variable is unclear in the case of Chile: Hanson's (2005) market potential function is a combination of rapidly changing variables associated with rapid but geographically irregular economic development. Rapid household income growth has been geographically widespread but unevenly distributed (Modrego and Berdegué, 2015), and so has been the case of rapidly growing wages (Beyer and Le Foulon, 2002; Paredes, 2013) and heterogeneously growing housing prices (López and Aroca, 2012). Moreover, if persistence were a potential drawback to the use of lagged levels of the variable, lagged differences, while resolving the persistence problem, would nevertheless be small with little variation and explanatory power. Instead, we find that the best instruments appear to be lagged *differences*, not levels, suggesting that the within-group persistence of the market potential variable is not a serious concern with respect to the use of the Arellano-Bond instrumenting procedure.

The instrumentation strategy was supported by the results of the well-known Hansen test of overidentifying restrictions and also by the "Difference-in-Hansen" specification test of exogeneity of subsets of instruments (see Arellano and Bond, 1991). The identifying assumption of absence of serial correlation of errors in the model in levels (equation (24)) was tested with the use of the Arellano and Bond (1991) second-order serial correlation test of errors in the differenced specification. These tests, along with consistency with parameter restrictions given by the theory, were the criteria to select the use of lagged levels of the dependent variable and lagged *differences* of the endogenous regressors as the best available instrumentation strategy.

Table 1 summarizes the regression results of the GMM estimation of the model (25). In columns (1) and (2), we present the results for the period 1990–2000, including observations for 143 comunas in the estimation sample. In columns (3) and (4), we present the results for the period 2000–2009, where the number of survey rounds falls from six to four, although the number of comunas in the estimation sample increases to 300, because of the inclusion of a large number of smaller municipalities not sampled during the 1990s. In columns (5) and (6), the results are for the whole 1990–2011 period, what yields 324 comunas in the estimation sample. Columns (1), (3), and (5) report the estimates for the base model, that is, without including additional regressors other than those implied by the relationship in equation (23). In columns (2), (4), and (6), we include the additional controls. Both the Hansen J and the Arellano-Bond serial correlation tests support model specification in our four main variants of model (25) in columns (1)–(4). The Moran I test fails to reject the null of absence of spatial correlation of errors in four out of six specifications. Comparing the results for the cases that reject and that do not reject, one cannot see qualitative differences in the point estimates or their standard errors as to suspect of biases or important efficiency problems.¹⁶

For the first estimation period, the base model (column 1) yields an estimated elasticity of self-employment rates with respect to market potential of 0.52, which is highly significant even with Windmeijer-corrected standard errors. This value is around twice as large as the nonlinear, two-step GMM estimates in Modrego et al. (2014) for the elasticity of the number of firms per capita with respect to the market potential measure (between 0.23 and 0.25). The estimated elasticity of the self-employment rate with respect to domestic wages (−0.78 in column 1 of Table 1) is in line with theoretical expectations and significant at conventional levels. Furthermore, restriction (26b)—that casts the elasticity of substitution as the negative of the ratio between the marginal effects of wages and market potential—yields a value of around 1.5, very close to the *a priori* value (1.6) used in the MP calculation in order to avoid the nonlinearity of the parameters in estimation. This comparison gives further support to our estimates, since the *a priori* value of σ was obtained for a model of the number of firms per capita, a metric of business activity closely related, but certainly different, to the dependent variable in this present analysis. Moreover, these previous estimates were based on a different dataset and a different estimation period (2005–2010).

Finally, the autoregressive term is positive and statistically significant. The estimated coefficient lays between the downward-biased fixed-effect estimation and the upward-biased OLS in levels, both unreported, which serves as an additional specification test (see Bond, 2002).¹⁷ The positive and significant coefficient for the lag dependent variable would confirm the importance of dynamic adjustments in local entrepreneurship rates. In particular, this result supports for the Chilean case the existence of a path-dependent local entrepreneurship dynamics, a pattern also documented for business start-up rates in Sweden (Andersson and Koster, 2011). Along with the fairly significant effect of our two main explanatory variables (wages and MP) the regressions results in column (1) point at spatiotemporal persistence of local entrepreneurship in the country (Fritsch and Mueller, 2007; Andersson and Koster, 2011).¹⁸

The inclusion of additional regional controls (column 2) brings no major qualitative changes in the results compared to those in column (1). Estimated MP and wage

¹⁶We are grateful to an anonymous reviewer for suggesting the application of this test.

¹⁷Available upon request

¹⁸The reader should be careful of not overinterpreting this parameter since there are could be many reasons simultaneously operating for this inertia, reasons we are not trying to explain here. We thank an anonymous reviewer for remarking this point.

TABLE 1: Estimation Results for Model (25).

Parameters (Standard Errors) ¹	(1)	(2)	(3)	(4)	(5)	(6)
Lag log entrep. rate	0.250*** (0.082)	0.159*** (0.079)	0.130 (0.083)	0.042 (0.077)	0.130*** (0.052)	0.066 (0.057)
Log Market Potential	0.515*** (0.205)	0.567*** (0.188)	0.295*** (0.140)	0.408*** (0.131)	0.311*** (0.109)	0.329*** (0.113)
Log wage	-0.782*** (0.348)	-0.763*** (0.308)	-0.404*** (0.197)	-0.435*** (0.181)	-0.428*** (0.154)	-0.425*** (0.151)
Log share with university studies		-0.078 (0.148)		-0.094 (0.060)		-0.041 (0.049)
Diversification index		-2.616*** (1.047)		0.057 (0.328)		0.068 (0.129)
Log share of women		0.350 (0.387)		-0.021 (0.286)		0.190 (0.238)
Log share in mid-labor career		-0.049 (0.129)		-0.136 (0.087)		-0.089 (0.067)
Implicit Parameter						
Sigma	1.517*** (0.402)	1.346*** (0.461)	1.370*** (0.391)	1.065*** (0.356)	1.375*** (0.266)	1.291*** (0.277)
Statistics						
N	361	361	576	574	1,440	1,435
Global significance P-value	0.003	0.030	0.040	0.015	0.000	0.000
P-value serial correlation test (1st or.)	0.000	0.000	0.000	0.000	0.000	0.000
P-value serial correlation test (2nd or.) ²	0.930	0.417	n.a.	n.a.	0.247	0.196
Instruments set ³	(1)	(2)	(1)	(2)	(1)	(2)
P-value Hansen/Sargan J	0.343	0.400	0.351	0.511	0.311	0.325
Moran's I of errors (P-value)	0.315	0.298	0.000	0.000	0.301	0.281
Estimation period	1990–2000	1990–2000	2000–2009	2000–2009	1990–2011	1990–2011

Notes: ¹All specifications include year dummies. Windmeijer-corrected standard errors. ²n.a. = not available. ³Instrument set (1): Second lag entrepreneurship rates (including employers and self-employed), lagged difference of the market potential function and lagged and twice-lagged differences of wages. Instrument set (2): Second lag entrepreneurship rates (including employers and self-employed), lagged difference of the market potential function, lagged and twice-lagged differences of wages, lagged difference of additional controls. Dependent variable: log rates of employers + self-employed in Chilean comunas.

elasticities are very similar and still significant, but the coefficient for the lagged entrepreneurship rates decreases substantially (0.16). The implicit elasticity of substitution (σ) remains above one (1.3), highly significant and with a 95 percent confidence interval containing the *a priori* value used in constructing the MP variable. Notably, additional controls provide no explanatory power, with the exception of the economic diversity index, showing a negative coefficient that suggests specialization (*Marshallian*) instead of diversification externalities. Still, the Wald test fails to reject the joint insignificance of these four additional variables ($P = 0.17$). Moreover, the sign of the education parameter, as well as those for the share of women and workers in mid-career are all contrary to the initial expectations, although highly not significant individually.

Results for the 2000–2009 period in columns (3) and (4) preserve the main conclusions obtained for the 1990s. First, the estimates confirms the significant positive and negative effect of the market potential variable and wages respectively, either with (column 3) or without (column 4) the inclusion of additional controls. The elasticity of entrepreneurship rates with respect to the market potential variable is now much closer to estimates by Modrego et al. (2014), which were obtained for a closer period (2005–2010). Comparing both specifications across periods (i.e., columns 1 with 3 and 2 with 4), the implicit elasticities of substitution are now slightly lower, around 1.3 and 1.1 respectively, but still highly significant and with a 95 percent confidence interval that contains the *a priori* value in both cases. Again, additional controls are not statistically significant, not even the diversification index. The most notable change is, nevertheless, the considerable reduction in the estimated coefficient for the lagged dependent variable, which is now not significant. This is to some extent an expected result, as persistence should decrease as the time interval between observations grows. Nevertheless, the sharp reduction and the loss of significance of the autoregressive coefficient in columns (3) and (4) is notable, because the length of the intervals for the 2000–2009 period is only one year longer than that of 1990–2000. Finally, although the Hansen J test supports the validity of the instruments, due to the small number of periods we were unable to perform the A-B test of second-order serial correlation, so we cannot fully dismiss the possibility of specification errors.

Finally, columns (5) and (6) report our (theoretically biased) estimates for the whole period. The main conclusion is that, given our main hypotheses, not considering the variable interval length for the data brought no serious consequence given the main hypotheses in this particular application. As expected, the autoregressive parameters lay in between those obtained for the two- and three-years intervals, which in this case implies inconclusive results in regard of entrepreneurship persistence. More important, for our main two variables, the expected sign and significance is preserved. Parameter restriction (26b) still holds and remains very similar (1.3–1.4) to the estimates for the other periods and to the *a priori* value. Finally, comparing the results in columns (1), (3), and (5) with those in columns (2), (4), and (6), one can only come to the conclusion that it is hard to anticipate the effect of omitting the varying-time interval in the case of the nonautoregressive covariates. As shown in Millimet and McDonough (2013), such effect will in general depend on their own dynamic behavior.

*Some Robustness Checks*¹⁹

The complexity of entrepreneurship as an economic phenomenon always makes the selection of an empirical metric a difficult task (see for instance Glaeser and Kerr, 2009;

¹⁹This subsection was largely motivated by the comments of one *JRS* reviewer to whom we are very grateful.

Parker, 2009). In this particular case, it could be argued that the share of business owners in a municipality could be a better empirical counterpart for the entrepreneurship rates in the theoretical model. Therefore, we re-estimate model (25) by restricting the dependent variable to a measure of the share of nonfarm employers only.

Table 2 summarizes the results obtained using this employer-only alternative dependent variable. Comparing the results in Tables 1 and 2, we conclude that the main relationships in this study are still robust to the redefinition of entrepreneurship. In particular, coefficients associated to the market potential variable remain positive and significant in the six model specifications, although the rate of employers proves to be more sensitive to changes in market potential than the rate as measure by employers plus self-employed. Similarly, the coefficient on wages remains larger than the elasticity with respect to the MP function in the six model specifications. Therefore, the implicit elasticity of substitution (26b) remains greater than one, and actually within the same range as those reported in Table 1. The most notable change in this set of estimates is the small and insignificant parameter for the autoregressive term, even for the 1990–2000 period. This last result indicates that there is little or no evidence of persistence in local entrepreneurship in Chile when measured by the rate of employers in the labor force.

Taking together, the results in Tables 1 and 2 are theoretically consistent and statistically robust with respect to model specification and estimation period. The results confirm the positive partial effect of the MP and the negative effect of domestic wages on municipal self-employment rates in the country. More importantly, in all the estimations the implicit elasticity of substitution is consistent with theoretical expectations (greater than one), and not substantially different, statistically speaking, from the previously estimated value for Chile obtained from a nonlinear model (Modrego et al., 2014).

Second, we performed further robustness checks by varying the *a priori* elasticity of substitution in the computation of the market potential variable. For values of $\sigma = [2, 3, 4]$, the six specifications in Tables 1 and 2 yielded partial elasticities of regional entrepreneurship rates with respect to market potential and wages according to model's predictions and individually significant at least at the 10 percent. Furthermore, the parameter restriction (Equation 26b) is verified and yielded intervals that at standard confidence levels contain the *a priori* value.²⁰

Third, modifiable aerial unit considerations are presumably important, as *comunas* could not be the most accurate representation of Chilean local economies. We re-estimated the model at the level of 44 provinces, which means with less than 200 and 100 observations for each subperiod respectively.²¹ Although in this case finite sample biases of the A-B estimator impedes being confident about the results (nor one can expect too many significant parameters with the Windmeijer correction), point estimates proved to be qualitatively similar to municipal estimates for the model with the base specification (columns 1, 3, and 5 of Table 1). In the case of the model of entrepreneurship measured as the share of employers only, in general, it also supports the model's implication for both with and without additional controls for the periods 1990–2011 and 2000–2009 (columns 3–6 of Table 2); although the base specification for the period 1990–2011 yielded a parameter restriction (26b) lower than one.

Fourth, although officially declared as representative for “self-represented comunas” (Mideplan, 2005), CASEN survey has been criticized for providing inaccurate estimates at the municipal level, particularly for smaller comunas (see for instance Agostini and Brown, 2011). For this reason, we ran the regressions by weighting observations by the

²⁰Available upon request.

²¹Available upon request.

TABLE 2: Estimation Results for Model (25).

Parameters (Standard Errors) ¹	(1)	(2)	(3)	(4)	(5)	(6)
Lag log entrep. rate	0.005 (0.065)	0.024 (0.067)	-0.010 (0.059)	0.006 (0.065)	0.037 (0.045)	0.014 (0.048)
Log market potential	2.136*** (0.688)	2.097*** (0.735)	0.846** (0.425)	0.975** (0.482)	0.915*** (0.277)	0.979*** (0.316)
Log wage	-3.544*** (1.089)	-3.275*** (1.082)	-1.164* (0.782)	-1.432* (0.782)	-1.227*** (0.435)	-1.044** (0.440)
Log share with university studies		0.015 (0.439)		-0.018 (0.312)		-0.033 (0.179)
Diversification index		-2.124 (3.443)		-1.374 (2.366)		-0.429 (0.488)
Log share of women		-0.955 (0.934)		0.043 (1.404)		0.350 (0.675)
Log share in mid-labor career		-0.126 (0.379)		0.277 (0.308)		0.077 (0.221)
Implicit Parameter						
Sigma	1.659*** (0.248)	1.561*** (0.321)	1.376*** (0.435)	1.469** (0.624)	1.341*** (0.293)	1.066*** (0.325)
Statistics						
N	327	327	479	479	1,179	1,179
Global significance P-value	0.001	0.007	0.000	0.000	0.000	0.000
P-value serial correlation test (1st or.)	0.001	0.002	0.000	0.000	0.000	0.000
P-value serial correlation test (2nd or.) ²	0.256	0.217	n.a.	n.a.	0.533	0.523
Instruments set ³	(3)	(4)	(3)	(4)	(3)	(4)
P-value Hansen/Sargan J	0.218	0.682	0.417	0.327	0.152	0.280
Moran's I of errors (P-value)	0.629	0.632	0.000	0.000	0.242	0.241
Estimation period	1990–2000	1990–2000	2000–2009	2000–2009	1990–2011	1990–2011

Notes: ¹All specifications include year dummies. Windmeijer-corrected standard errors. ²n.a. = not available. ³Instrument set (3): Second lag entrepreneurship rates (including employers only), lagged and twice-lagged differences of the market potential function and lagged and twice-lagged differences of wages. Instrument set (4): Second lag entrepreneurship rates (including employers only), lagged and twice-lagged differences of the market potential function, lagged and twice-lagged difference of wages, lagged difference of additional controls. Dependent variable: log rates of employers in Chilean comunas.

inverse of the number of households surveyed in each comuna in each CASEN round.²² Results for the two subperiod remains the same. For the whole period, estimates equivalent to columns 5 and 6 of Table 1 and column 6 of Table 2 delivered implicit elasticities of substitution lower than 1 (although close).

Fifth, to analyze the potential role of the particular geography of extreme regions of the country, we ran the regressions excluding comunas in the regions of Arica and Parinacota, Antofagasta and Atacama (extreme north) and Aysén and Magallanes (extreme South). As correctly observed by one of the reviewers, these are low-density areas with comunas that are extremely separated from each other, and where the MP could well be given by the comuna itself. Results did not change in any significant way compared to the full-sample models.²³

Finally, since our empirical proxy of municipal MP's does not consider own incomes for all but 51 comunas sampled during the whole 1990–2011 period, we re-estimated the model with only these 51 comunas. Due to the small sample size, we conducted this exercise only for the 1990–2011 period. Results remained qualitatively unaltered.²⁴

Taken as a whole, the empirical evidence provides a consistent support to the location-driven trade-off between profit effects and opportunity-costs effects conditioning the regional supply of entrepreneurs in Chile. Consequently, the spatial occupational choice model proposed here provides a convincing explanation to the weak spatial association between market potential and entrepreneurship rates in Chile, and perhaps for other developing countries showing a similar pattern. One may argue that given the large value used for the distance decay parameter in the computation of the market potential variable, our results could be observationally equivalent to the effects of domestic market size and/or urbanization externalities in domestic markets. We believe that this is not the case here, as the market potential was built taking into account the distance to 51 major municipalities, and therefore, for most observations in the sample, home-market effects are not included.

6. CONCLUSION

The development of spatial economic models and methods has increased our understanding of the role of spatial demand linkages and localization externalities in shaping the geographic distribution of entrepreneurial activity. To date, however, most studies have fallen short in crafting microeconomic foundations for the role of location in the individual's decision to become an entrepreneur. This paper aims at contributing to such a line of research.

The proposed model is one of the few NEG-based framework that makes explicit the trade-offs driven by market potential in the occupational choice of entrepreneurship versus wage employment. Higher MP has a positive or pull effect on entrepreneurship through providing market opportunities for business profits, but at the same time, it has a negative influence through its positive effect on wages, the opportunity cost of being an entrepreneur. In the long run, differences in entrepreneurship rates across regions, are given by differences in average fixed costs across firms, which in turn depend on differences in the distribution of entrepreneurial ability. The effects of economic geography on entrepreneurship are thus ultimately exerted by shaping migration patterns across

²² Available upon request.

²³ Available upon request.

²⁴ Available upon request.

regions. Overall, sorting matters (Combes et al., 2008) and models of agglomeration with heterogeneous agents (see for instance Behrens and Robert-Nicoud, 2015) are a promising avenue to build structured hypotheses regarding the role of regional market potential on the location of entrepreneurial individuals.

The model presented in this paper makes a twofold contribution. For the spatial economics literature, it introduces explicitly entrepreneurs' heterogeneity as another factor that, along with agglomeration externalities, can determine regional differences in entrepreneurship rates, average firm size, average firm costs, and labor productivity. The implication is that *everything else equal*, locations with better market access allow less cost-efficient (able) firms (entrepreneurs) to thrive. For the occupational choice literature, the model addresses the effect of location on the two key economic variables that determine the selection into entrepreneurship: entrepreneurial profits and wages.

The model proposed here inherits many of the rigidities of the standard spatial Dixit-Stiglitz framework, especially its structure of monopolistic competition with constant mark-up over marginal labor costs. It is this structure what leads to the result that, in equilibrium, the profit and opportunity cost-effect exactly offset each other. But at the same time, it is flexible enough to allow different equilibrium results. For instance, one can relax some simplifying assumptions, such as managerial ability affecting fixed costs only, or one can introduce additional elements, such as nonpecuniary motivations to choose entrepreneurship or spatial differences in start-up costs. Such extensions could well account for different geographic patterns perhaps observed in other countries in spatially-explicit occupational choice settings.

In the case of Chile, our simplifying assumptions allow us to propose an econometric specification suitable for testing the countervailing effects of market potential on entrepreneurship with the limited data at hand, which is our main motivation given the weak spatial association between these variables in Chile. The results strongly support the formal mechanics of the model, at least as a long-run equilibrium relationship towards which Chilean local economies move. Further evidence seems necessary to establish a broader empirical validation of the model in other national contexts, and in particular, to test for potential differences between developed and developing countries, where drivers of local entrepreneurship likely differ.

Finally, this paper has left open some interesting questions that deserve more theoretical work as well as empirical evidence. For instance, the model's equilibrium results would likely differ in conditions where long-run interregional migration flows might be restricted. It would be interesting to explore the long-run equilibrium with more explicit treatment of migration costs, allowing a correlation between such cost and the individual's entrepreneurial ability. This is a relevant question, because migrants are potentially different in their managerial skills and are perhaps differentially able to take advantage of place-specific assets and resources. Also, it would be important to extend the model in order to incorporate the role of amenities in explaining the spatial variation of entrepreneurship, as it has been done in some new NEG models (Sudekum, 2008) and urban economic models of entrepreneurship (Glaeser, Kerr, and Ponzetto, 2010; Glaeser, Rosenthal, and Strange, 2010).²⁵ Finally, it is possible to incorporate and test for different mechanisms behind the spatial variation of entrepreneurship induced both by opportunity and by necessity (e.g., Stephens et al., 2013).

²⁵We thank an anonymous reviewer for suggesting this extension.

APPENDIX 1.

PROOF OF HYPOTHESES 1 AND 2

Totally differentiating (7) and rearranging terms yields

$$(A.1) \quad W \frac{dz_r}{dz_r} + \sigma \frac{dw_r}{w_r} = \frac{dMP_r}{MP_r},$$

with $W = -\gamma \frac{(\phi MP_r w_r^{-\sigma} - 1)}{\phi MP_r w_r^{-\sigma}} < 0$, since the marginal entrepreneur makes nonnegative profits.

And differentiating (11) and rearranging terms yields

$$(A.2) \quad \sigma \frac{dw_r}{w_r} + Z \frac{dz_r}{z_r} = \frac{dMP_r}{MP_r},$$

with $Z = \frac{w_r^\sigma z_r}{\omega MP_r} [\frac{F'(z_r)}{(1-F(z_r))^2} - \bar{k}'(z_r)] > 0$, since $F'(z_r) > 0$ and $\bar{k}'(z_r) < 0$.

Stacking (A.1) and (A.2) in a displacement matrix:

$$\begin{bmatrix} W & \sigma \\ Z & \sigma \end{bmatrix} \begin{bmatrix} d \ln z_r \\ d \ln w_r \end{bmatrix} = \begin{bmatrix} 1 \\ 1 \end{bmatrix} d \ln MP_r.$$

Rearranging gives:

$$\begin{bmatrix} \frac{d \ln z_r}{d \ln MP_r} \\ \frac{d \ln w_r}{d \ln MP_r} \end{bmatrix} = \begin{bmatrix} W & \sigma \\ Z & \sigma \end{bmatrix}^{-1} \begin{bmatrix} 1 \\ 1 \end{bmatrix} = \Omega^{-1} \begin{bmatrix} 1 \\ 1 \end{bmatrix},$$

$\Omega^{-1} = \frac{1}{|\Omega|} \begin{bmatrix} \sigma & -\sigma \\ -Z & W \end{bmatrix}$, which implies $\frac{d \ln z_r}{d \ln MP_r} = 0$, which proves hypothesis 2.

In addition, $|\Omega| = \sigma(W - Z)$, so

$$(A.3) \quad \frac{d \ln w_r}{d \ln MP_r} = \frac{1}{\sigma},$$

which proves hypothesis 1.

APPENDIX 2.

WELFARE ANALYSIS OF THE SHORT-RUN EQUILIBRIUM

Differentiating the profit equation (5), using (A.3) and substituting terms:

$$(B.1) \quad \frac{d\pi_{ir}}{dMP_r} = \frac{\phi MP_r w_r^{1-\sigma}}{MP_r} + \frac{(1-\sigma)\phi MP_r w_r^{1-\sigma} - w_r \tau \theta_i^{-\gamma}}{w_r} \frac{dw_r}{dMP_r} = \frac{\pi_{ir}}{MP_r} \frac{1}{\sigma}.$$

Equations (A.3) and (B.1) mean that all participants gain proportionally the same from an increase in the regional MP: $\frac{d \ln(\pi_{ir})}{d \ln(MP_r)} = \frac{d \ln(w_r)}{d \ln(MP_r)} = \frac{1}{\sigma}$.

To analyze the allocation of surplus in aggregate terms, let Π_r be the total surplus in the region which comprises the sum of total entrepreneurial (Π_r^e) and total wage (Π_r^w) rents:

$$\Pi_r = L_r \left(\int_0^{z_r} w_r f(\theta) d\theta + \int_{z_r}^{+\infty} \pi_{ir} f(\theta) d\theta \right) = \Pi_r^w + \Pi_r^e.$$

Using (A.3) and (B.1), the effect of a shock in the regional market potential on the total surplus is $\frac{d\Pi_r}{dMP_r} = \frac{w_r}{\sigma MP_r} l_r + \frac{\bar{\pi}_r}{\sigma MP_r} e_r$, where $\bar{\pi}_r$ indicates average business profits.

The relative additional rents triggered by the increase in the market potential are

$$\frac{\left(\frac{d\pi_c}{dMP_r}\right)}{\left(\frac{d\pi_w}{dMP_r}\right)} = \frac{\bar{\pi}_r e_r}{w_r \bar{l}_r} = \frac{\bar{\pi}_r}{\pi(z_r)} \frac{e_r}{\bar{l}_r},$$

that is, the relative gains of both groups depend only on their relative sizes.

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