

The Spillover Effects of Top Income Inequality*

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Abstract

Since the 1980s top income inequality has increased considerably within occupations as diverse as bankers, managers, doctors, lawyers and scientists. Such a broad pattern has led the literature to search for a common explanation. We show instead that increases in income inequality originating within a few occupations can “spill over” into others, driving broader changes in income inequality. We develop an assignment model where *generalists* with heterogeneous income buy services from *doctors* with heterogeneous ability. In equilibrium the highest-earning generalists match with the highest quality doctors. Increases in income inequality among the generalists feed directly into the doctors’ income inequality. To test our theory, we identify occupations for which our consumption-driven theory predicts spillovers and occupations for which it does not. Using a Bartik-style instrument, we show that an increase in general income inequality causes higher

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income inequality for doctors, dentists and real estate agents; and in fact accounts for most of the increases in inequality within these occupations. Physician pricing and insurance network data support our mechanism.

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1 Introduction

Since the 1980s the share of total earnings going to the top of the income distribution has increased considerably. At the same time income inequality within the top has also increased with a higher share of top earnings going to the very high earners. Moreover this pattern holds within high-earning occupations so that the overall growth of top income inequality is not simply due to the growth within particular occupations (Bakija, Cole, and Heim, 2012). At first glance, this broad pattern suggests that a plausible explanation—whether it be globalization, deregulation, changes to the tax structure, or technological change—would have to apply to occupations as diverse as financial managers, doctors, and CEOs (Kaplan and Rauh, 2013). We argue that this need not be the case because inequality across occupations is linked. We show that exogenous increases in income inequality within one occupation “spill over” into others through the former’s consumption, driving up income inequality for a broader set of occupations than those affected by the initial shock.

We present a model where changes in within-occupation income inequality propagate to other occupations through consumption, rather than through competition for skill in the broader labor market. We study an assignment model where *generalists* with heterogeneous income buy the services of *doctors* with heterogeneous ability. In equilibrium the highest-earning generalists match with the highest-ability doctors. Increases in income inequality among the generalists feed directly into income inequality among doctors. Two conditions on the services provided by doctors are necessary for the equilibrium to feature an assignment mechanism and thereby income inequality spillovers: *heterogeneity* and *non-divisibility* in output. Non-divisibility means that one high-ability doctor is not the same as two decent-ability doctors. We focus on physicians, dentists and real estate agents, occupations that meet these conditions, and contrast them with occupations that do not. Using data from the Decennial Census and the American Community Survey, we find that an increase in general income inequality causes an increase in inequality for these occupations, with a spillover elasticity ranging from 0.5 to 2.7. These occupations are important within the top 1%—in fact Physicians are the most common Census occupation in the top 1% in 2014.

Our baseline model considers occupations of heterogeneous ability where production is not scalable; that is, no mechanism exists that would allow the more talented to scale up output. For these occupations, when consumption is non-divisible, the income distribution is tightly linked to that of the general population. Specifically, generalists of

heterogeneous ability produce a homogeneous product in quantity proportional to their skill level. Besides this homogeneous product, each generalist consumes the services of one doctor. Doctors also have heterogeneous ability but their ability translates proportionately into the quality of the services they provide and not the quantity. All doctors serve the same number of patients. The abilities of both generalists and doctors are Pareto distributed but with different parameters. The result is an assignment function with positive assortative matching: the highest ability generalists match with the highest ability doctors. An exogenous mean-preserving spread in the income inequality of generalists increases the number of high-earning generalists, increases the demand for the best doctors and increases top income inequality among doctors as well. In fact, in the special case of Cobb-Douglas utility, top income inequality of doctors is entirely driven by the earnings distribution of generalists and is independent of the underlying ability distribution of doctors.

We extend the model in three directions. First, we allow for occupational mobility at the top: high-ability doctors can choose to be high-ability generalists and vice versa. Since changes in top income inequality for generalists completely translate into changes in top income inequality for doctors when there is no occupational mobility, allowing doctors to switch occupation or not has no impact on doctors' inequality and the two settings are observationally equivalent. Second, we consider two regions in one nation that differ only in the top income inequality of generalists and allow patients to import their medical services. We show that top income inequality among doctors for each region must follow generalist top income for the most unequal region. This distinction will be important for the empirical test: When a service is local, so non-tradable, spillover effects will happen at the local level, whereas for tradable services the spillover effect will happen at the national level. Finally, we let doctors move across regions and show that the most unequal region will attract the most able doctors, but, as in the baseline model, doctors' inequality is determined by general inequality in the region where they eventually live. Hence the observed top income inequality of doctors is the same whether or not they can move.

To test our model we take as a starting point the fact that top income inequality has increased broadly across occupations. The solid line of Figure 1 shows that the relative income of those in the top 0.1% relative to the top 1% of the income distribution has risen from 3.1 in 1979 to 4.3 in 2005. The pattern is similar for occupations as diverse as doctors, real estate agents, and scientists. It also holds for a number of occupations with

incomes mostly below the top 1% such as college professors and secretaries.¹ We test our theory using a combination of the Decennial Census and the American Community Survey for every decade since 1980 and we focus on labor market areas (an aggregation of commuting zones; Dorn, 2009) as the unit of analysis.² We construct measures of top income inequality that are specific to the year/occupation/labor market area, but since the income data are top-coded for around 0.5 per cent of observations — and therefore significantly more for high-earning occupations — we impose an assumption of a Pareto distributed right tail of the income distribution and use the exponential parameter of the estimated Pareto distribution as an (inverse) measure of income inequality. Our theory predicts which occupations will feature spillovers: those with non-divisibility in output. Furthermore, since we will focus on geographical variation across the United States, our estimation methodology will only pick up spillover effects if they are local—that is, if workers mostly service local clients. We classify occupations into two groups: Those that meet these conditions (such as physicians, dentists and real estate agents) and those that do not (such as financial managers, college professors and secretaries.) Using panel data, in OLS regressions, we find that an increase in general income inequality (excluding the occupation of interest) is positively correlated with an increase in inequality for occupations in the first group, but mixed results for the other occupations. Naturally state-specific changes in regulation, labor demand, or taxes might cause occupation-specific income inequality to increase at the same time as general income inequality. To establish a causal link, we use a Bartik (1991)-style instrument. We construct a weighted average of nationwide inequality for the 20 occupations that are the most represented in the top 5% nationwide (excluding the occupation of interest). The weights correspond to the relative importance of each occupation in each labor market area at the beginning of our sample. In other words, we only exploit the changes in labor market income inequality that arise from the occupational distribution in 1980 combined with the nationwide trends in occupation-specific inequality. This weighted average serves as our instrument for general inequality in the area in question.

Using this instrument, we find a very clear distinction between the first group and other occupations. An increase in general income inequality at the local level causes an increase in inequality for physicians, dentists and real estate agents, who operate in local

¹This is not a result of large changes in the occupational distribution in the top 1% and top 0.1%. Except for financial professionals, whose weight in the top has increased substantially, the distribution of the largest occupations in the top has remained relatively constant from 1979 to 2005.

²At the moment we use the publicly available data provided by IPUMS (Ruggles et al., 2015). We are in the process of obtaining access to the uncensored data.

markets. The parameter estimates are consistent with the majority of the increase in income inequality for these occupations being explained by increases in general income inequality. On the other hand, as our theory predicts, we find that *local* general income inequality does *not* spill over to financial managers, college professors and secretaries.

Our model proposes a specific mechanism for transmitting inequality from the general population to private physicians: price inequality and assortative matching. Physicians in more unequal areas should charge unequal prices. We use detailed physician claims data from three states to directly examine this mechanism. Since actual physician payments in the United States reflect the structure of insurance plan networks, we also examine inequality in these network sizes. Both types of data support the mechanism we propose: pricing inequality is increasing in areas with growing inequality, and network size is more heterogeneous in more unequal markets.

The increase in top income inequality has inspired substantial scholarship (among many others, see Piketty and Saez, 2003, and Atkinson, Piketty and Saez, 2011). This literature has established that at the top, the income distribution is well-described by a Pareto distribution (see Guvenen, Karahan, Ozkan, and Song, 2015, for some of the most recent evidence, and Pareto, 1896, for the earliest). Further, Jones and Kim (2014) show that the increase in top income inequality is linked with a fattening of the right tail of the income distribution, which corresponds to a decrease in the shape parameter of the Pareto distribution. This literature is related to, but distinct from, the large literature on skill-biased technological change and income inequality which seeks to explain changes in income inequality throughout the income distribution and primarily *across* occupations (Goldin and Katz, 2010; Acemoglu and Autor, 2011.)

More specifically, our paper builds on the “superstars” literature originating with Rosen (1981), who explains how small differences in talent may lead to large differences in income. The key element in his model is an indivisibility of consumption result which arises from a fixed cost in consumption per unit of quantity. This leads to a “many-to-one” assignment problem as each consumer only consumes from one performer (singer, comedian, etc.), but each performer can serve a large market (see also Sattinger, 1993).³ In that framework, income inequality among performers increases because technological change or globalization allows the superstars to serve a much larger market—that is, to scale up production. Specifically, if $w(z)$ denotes the income of an individual of talent z , $p(z)$ denotes the average price for his services, and $q(z)$ is the quantity provided,

³Adding network effects, Alder (1985) goes further and writes a model where income can drastically differ among artists of equal talents.

such that $w(z) = p(z)q(z)$, the standard interpretation of “superstars” is that they have very large markets (a high $q(z)$). This makes such a framework poorly suited for occupations where output is not easily scalable, such as doctors. In contrast we focus on such occupations and study an assignment model that is “one-to-one” (or more accurately “a constant-to-one”) where superstars are characterized by a large price for their services $p(z)$. This makes our paper closer to Gabaix and Landier (2008) who build a “one-to-one” assignment model to study CEOs’ compensation. They argue that since executives’ talent increases the overall productivity of firms, the best CEOs are assigned to the largest firms, and show empirically that the increase in CEO compensation can be fully attributed to the increase in firms’ market size (Grossman (2007) builds a model with similar results). Along the same lines, Määttänen and Terviö (2014) build an assignment model to study house price dispersion and income inequality. They calibrate their model to six US metropolitan areas and find that the increase in inequality has led to an increase in house price dispersion. Gabaix, Lasry, Lions and Moll (2015) argue that the fast rise in both the share of income held by the top earners and income inequality among these earners requires aggregate shocks to the return of high income earners (“superstar shocks”). Our analysis suggests that even if such shocks only directly affect some occupations they will spill over into other occupations. The original shock may arise from technological change in occupations where span-of-control features are pervasive as suggested by Geerolf (2015).⁴ Globalization can increase the share of income going to the top earners and also increase inequality among these earners (see Manasse and Turini, 2001; Kukharskyy, 2012; Gesbach and Schmutzel, 2014 and Ma, 2015⁵).

Beyond “superstar” effects, the economics literature has investigated several possible explanations for the rise in top income inequality. Regardless of what the underlying shock or shocks may be, our paper shows how it can spill over into the broader economy. These spillovers could make it difficult to test between the various hypotheses that have been proposed about the key underlying shock. Jones and Kim (2014) and Aghion, Akcigit, Bergeaud, Blundell and Hémous (2015) look at the role played

⁴Geerolf (2015) builds a span of control model to micro-found the fact that firms’ size distribution follows Zipf’s law. His model naturally leads to “superstar” effects and a bounded distribution of talents can lead to an unbounded distribution of income. Similarly, Garicano and Hubbard (2012) build a span of control model which features positive assortative matching as the most skilled individuals become the most skilled managers who manage large firms which employ the most skilled workers. They use data from the 1992 Census of services on law offices to find support for their model. Yet, span of control issues do not seem *directly* relevant for doctors or real estate agents.

⁵Yet, none of these papers are able to generate a change in the shape parameter of the Pareto distribution of top incomes through globalization.

by innovation;⁶ Piketty (2014) argues that top income inequality has increased because of the high returns on capital that a concentrated class of capitalists enjoy; Piketty, Saez and Stantcheva (2014) argue that low marginal income tax rates divert managers' compensation from perks to wages and increase their incentive to bargain for higher wages. Philippon and Reshef (2012) emphasize the role played by the financial sector, and Böhm, Metzger and Strömberg (2015) question whether this premium reflects true talent. For our purposes, all that matters is that consumers have heterogeneous preferences, whether or not those preferences reflect realtors' actual marketing skills or surgeons' actual cutting skills.

We proceed to present the theoretical model in Section 2. Section 3 contains our empirical strategy, data, and description of the instrument. Section 4 includes our core empirical results. In section 5 we use data on physician networks and pricing to directly test the mechanism embedded in the model. We conclude in Section 6.

2 Theory

We first present our baseline model. We consider occupations of heterogeneous ability where production is not scalable—there is no mechanism that would allow the more talented to scale up output—and consumption is non-divisible. In this case, we demonstrate that the within-occupation income distribution is tightly linked to that of the general population. To help guide our empirical analysis and determine when we would expect to see spillover effects in the data, we then relax a number of assumptions. “Doctors” will represent occupations where the most skilled workers can produce a good of higher quality but cannot serve more customers than the less skilled, and where customers cannot divide their consumption across several producers. One high-ability doctor is not the same as two decent-ability doctors. Besides doctors, prominent examples are dentists, college professors, and real estate agents.

⁶Jones and Kim (2014) build a model close to the superstars literature where the distribution of income for top earners is Pareto and results from two forces: the efforts of incumbents to increase their market share and the innovations of entrants who can replace incumbents. Using a panel analysis of US states, Aghion et al. (2015) show empirically that an increase in innovation leads to more top income inequality.

2.1 The Baseline Model

We consider an economy populated by two types of agents: generalists of mass 1 and (potential) doctors of mass μ_d .

Production. Generalists produce a homogeneous good that serves as the numeraire. They differ in their ability to produce such that a generalist of ability x can produce x units of the homogeneous good. The ability distribution is Pareto such that a generalist is of ability $X > x$ with probability:

$$P(X > x) = \left(\frac{x_{\min}}{x}\right)^{\alpha_x},$$

with lower bound $x_{\min} = \frac{\alpha_x - 1}{\alpha_x} \hat{x}$ and shape $\alpha_x > 1$, which keeps the mean fixed at \hat{x} when α_x changes. The parameter α_x is an (inverse) measure of the spread of abilities. We will keep α_x exogenous throughout and will capture a general increase in top income inequality by a reduction in α_x . Doctors produce health services and can each serve λ customers, where we impose $\lambda \geq \max(1, 1/\mu_d)$ so that there are enough doctors to serve everyone. Potential doctors differ in their ability z , according to a Pareto distribution with shape α_z such that they will have ability $Z > z$ with probability:

$$P(Z > z) = \left(\frac{z_{\min}}{z}\right)^{\alpha_z}.$$

All potential doctors can alternatively work as generalist and produce the homogeneous good with ability x_{\min} (see section 2.3 for a model where doctors' and generalists' abilities are perfectly correlated). Though the ability of a doctor does not change how many patients she can take care of, it increases the utility benefit that patients get from the health services that are provided.

Consumption. Generalists consume the two goods according to the Cobb-Douglas utility function

$$u(z, c) = z^{\beta_z} c^{1-\beta_z}, \tag{1}$$

where c is the consumption of homogeneous good and z is the quality of the health care (equal to the ability of the doctor providing it).⁷ The notion that medical services are not divisible is captured by the assumption that each generalist needs to consume the services of exactly one doctor. This implies that there will not exist a common

⁷For our purpose, one should think of z as the quality of health care perceived by the consumers at the time when they decide on a doctor. So a pediatrician who can assuage an anxious parent might have a higher z than one with better diagnostic skills but fewer interpersonal skills.

price per unit of quality-adjusted medical services. For simplicity, doctors only consume the homogeneous good, an assumption that can easily be generalized (see section 2.5.1 below).

2.1.1 Equilibrium

Generalists. Since a generalist of ability x produces x units of the consumption good, their income must be distributed like their ability. The consumption problem of generalist of ability x can then be written as:

$$\begin{aligned} \max_{z,c} u(z,c) &= z^{\beta_z} c^{1-\beta_z}, \\ \text{st } \omega(z) + c &\leq x, \end{aligned} \tag{2}$$

where $\omega(z)$ is the price of one unit of medical services by a doctor of ability z .

Taking first order conditions with respect to the quality of the health services consumed and the homogeneous good gives:

$$\omega'(z) z = \frac{\beta_z}{1-\beta_z} (x - \omega(z)). \tag{3}$$

Since no generalist spends all her income on health care, this equation immediately implies that in equilibrium, $\omega(z)$ must be increasing such that doctors of higher ability earn more per unit of medical services. Importantly, the non-divisibility of medical services implies that doctors are “local monopolists” in that they are in direct competition only with the doctors of slightly higher or lower ability. As a consequence, doctors do not take prices as given, which implies that $\omega(z)$ will in general not be a linear function of z .

As a result, the equilibrium involves positive assortative matching between generalists’ income and doctors’ ability. We denote by $m(z)$ the matching function such that a doctor of ability z will be hired by a generalist whose income is $x = m(z)$ and $m(z)$ is an increasing function (see Appendix A.1 for a proof).

Doctors. Since there are (weakly) more doctors than needed the least able doctors will choose to work as generalists. We denote by z_c the ability level of the least able doctor who decides to provide health services so that $m(z)$ is defined over $[z_c, \infty)$ and $m(z_c) = x_{\min}$ (the worst doctor is hired by a generalist with income x_{\min}). Then, market

clearing at all quality levels implies that

$$P(X > m(z)) = \lambda \mu_d P(Z > z), \quad \forall z \geq z_c \quad (4)$$

There are $\mu_d P(Z > z)$ doctors with an ability higher than z , each of these doctors can serve λ patients, and there are $P(X > m(z))$ patients whose income is higher than $m(z)$. If $\lambda \mu_d > 1$ then $z_c > z_{\min}$ and if $\lambda \mu_d = 1$ then $z_c = z_{\min}$.

Using the assumption that abilities are Pareto distributed, we can use (4) to obtain the matching function as:

$$m(z) = x_{\min} (\lambda \mu_d)^{-\frac{1}{\alpha_x}} \left(\frac{z}{z_{\min}} \right)^{\frac{\alpha_z}{\alpha_x}}. \quad (5)$$

Intuitively if $\alpha_z > \alpha_x$, so that top talent is ‘scarcer’ among doctors than generalists then the matching function is convex because it must assign increasingly relatively productive generalists to doctors. At $m(z_c) = x_{\min}$, we obtain the ability of the least able potential doctor working as a doctor: $z_c = (\lambda \mu_d)^{\frac{1}{\alpha_z}} z_{\min}$. This is independent of the generalists’ income distribution because it only depends on quantities.

We denote by $w(z)$ the income of a doctor of ability z and note that $w(z) = \lambda \omega(z)$ since each doctor provides λ units of health services. Furthermore as a potential doctor of ability z_c is indifferent between working as a doctor and in the homogeneous good sector earning a wage equal to x_{\min} , we must have $w(z_c) = x_{\min}$. Now plugging the matching function in (3), we obtain the following differential equation which must be satisfied by the wage function $w(z)$:

$$w'(z) z + \frac{\beta_z}{1 - \beta_z} w(z) = \frac{\beta_z}{1 - \beta_z} x_{\min} \left(\frac{\lambda^{\alpha_x - 1}}{\mu_d} \right)^{\frac{1}{\alpha_x}} \left(\frac{z}{z_{\min}} \right)^{\frac{\alpha_z}{\alpha_x}}. \quad (6)$$

Using the boundary condition at $z = z_c$, we obtain a single solution for the wage profile of doctors which obeys (see Appendix A.2):

$$w(z) = x_{\min} \left[\frac{\lambda \beta_z \alpha_x}{\alpha_z (1 - \beta_z) + \beta_z \alpha_x} \left(\frac{z}{z_c} \right)^{\frac{\alpha_z}{\alpha_x}} + \frac{\alpha_z (1 - \beta_z) + \beta_z \alpha_x (1 - \lambda)}{\alpha_z (1 - \beta_z) + \beta_z \alpha_x} \left(\frac{z_c}{z} \right)^{\frac{\beta_z}{1 - \beta_z}} \right]. \quad (7)$$

One can show that $w(z)$ is increasing in doctor’s ability z as expected, with $w(z_c) = x_{\min}$. Intuitively, equation (7) consists of two parts: The first term dominates for large z/z_c and ensures an asymptotic Pareto distribution, and the second term fulfills the

indifference condition for the least able active doctor. Hence, for z/z_c large, we get that

$$w(z) \approx x_{\min} \frac{\lambda \beta_z \alpha_x}{\alpha_z (1 - \beta_z) + \beta_z \alpha_x} \left(\frac{z}{z_c} \right)^{\frac{\alpha_z}{\alpha_x}}. \quad (8)$$

Therefore, the wage schedule must be convex in z if $\alpha_z > \alpha_x$. To understand the intuition, consider again the case where top-talented doctors are scarce ($\alpha_z > \alpha_x$). This implies a fatter tail among generalists than doctors, such that a generalist of twice the income does not have a doctor of twice the ability. Hence, a linear schedule $\omega(z) \propto z$ cannot be an equilibrium as the Cobb-Douglas utility function would require a constant share spent on medical services, which would imply double the payment to a doctor that is not twice as good. For the same reason, the schedule cannot be concave when $\alpha_z > \alpha_x$.

We define $P_{doc}(W_d > w_d)$ as the probability that the wage of an actual doctor is higher than w (that is we only take into account the potential doctors who actually choose to work as doctors). We get that $P_{doc}(W_d > w_d) = (z_c/w^{-1}(w_d))^{\alpha_z}$, so that using (8), for w_d large enough:

$$P_{doc}(W_d > w_d) \approx \left(\frac{x_{\min} \lambda \beta_z \alpha_x}{\alpha_z (1 - \beta_z) + \beta_z \alpha_x} \frac{1}{w_d} \right)^{\alpha_x}. \quad (9)$$

That is, the income of (actual) doctors is distributed in a Pareto fashion at the top, with a shape parameter inherited from the generalists, independent of the spread of doctor ability, α_z . Similarly, the income distribution of potential doctors (denoted $P_{pot.doc}$) must then obey for w_d large enough:

$$P_{pot.doc}(W_d > w_d) \approx \frac{1}{\lambda \mu_d} \left(\frac{x_{\min} \lambda \beta_z \alpha_x}{\alpha_z (1 - \beta_z) + \beta_z \alpha_x} \frac{1}{w_d} \right)^{\alpha_x}.$$

In particular, a decrease in α_x directly translates into a decrease in the Pareto parameter for doctors' income distribution: an increase in inequality among generalists leads to an increase in inequality among doctors. In other words, the increase in top income inequality spills over from one occupation (the generalists) to another (doctors). At the top it also increases the income of doctors—as a decrease in α_x leads to an increase in $P(W_d > w_d)$ for w_d high enough.⁸ Formally:

⁸Not all doctors benefit, though, as we combine a decrease in α_x with a decrease in x_{\min} to keep the mean constant. As a result the least able active doctor, whose income is x_{\min} , sees a decrease in her income. Had we kept x_{\min} constant so that a decrease in α_x also increases the average generalist

Proposition 1. *Doctors' incomes are asymptotically Pareto distributed with the same shape parameter as the generalists'. In particular an increase in top income inequality for generalists increases top income inequality for doctors.*

Further, a decrease in the mass of potential doctors μ_d (or an increase in the mass of generalists, which we have normalized to 1 here) does not affect inequality among doctors at the top but it increases the share of doctors who are active (z_c decreases) and their wages (as $w(z)$ increases if z_c decreases).

Taking stock. Proposition 1 establishes the central theoretical result of our paper. For the empirical analysis it is important to establish which assumptions are necessary for the spillover result and which are not. We will do this in subsequent sub-sections. Before pursuing these extensions, however, we consider the implications of our basic result for other key outcomes in this market: health expenditures and welfare inequality.

2.1.2 Implications for spending and welfare

Health expenditures. Health care prices increase sharply at the top, in fact, thanks to the Cobb-Douglas assumption, we obtain that rich generalists spend close to a constant fraction of their income on health. Formally, a generalist with income x spends $w(m^{-1}(x))/\lambda$ in health services. Using (5), we obtain that his health spending of $h(x)$ must obey:

$$h(x) = \frac{\beta_z \alpha_x}{\alpha_z(1 - \beta_z) + \beta_z \alpha_x} x + \frac{1}{\lambda} \frac{\alpha_z(1 - \beta_z) + \beta_z \alpha_x(1 - \lambda)}{\alpha_z(1 - \beta_z) + \beta_z \alpha_x} x_{\min} \left(\frac{x}{x_{\min}} \right)^{-\frac{\alpha_x}{\alpha_z} \frac{\beta_z}{1 - \beta_z}}. \quad (10)$$

Note that health care is a necessity if $\alpha_z(1 - \beta_z) + \beta_z \alpha_x(1 - \lambda) > 0$. This follows from the price gradient consumers face (equation 7), which follows from the prices low-quality and high-quality doctors charge. Low-quality prices are pinned down by the indifference condition for the lowest doctor z_c . High-quality prices are determined purely by the parameters of the utility function and the ability distributions. Specially, consider the case in which $\alpha_z(1 - \beta_z) + \beta_z \alpha_x(1 - \lambda) > 0$: a doctor in the right end of the tail serving a patient of income x earns $\lambda \beta_z \alpha_x / (\alpha_z(1 - \beta_z) + \beta_z \alpha_x) x$. If the lowest quality doctor were to charge the same, she would earn $\lambda \beta_z \alpha_x / (\alpha_z(1 - \beta_z) + \beta_z \alpha_x) x_{\min} < x_{\min}$, which would be insufficient to compensate her for her outside option as a generalist earning x_{\min} . Consequently, she must be charging a larger share of patient income, and since

income, then all doctors would have benefited.

everybody consumes the services of exactly one doctor, medical services are a necessity. This is more likely to be the case when the number of patients a doctor can treat, λ , is low or when $\alpha_z > \alpha_x$ so generalists have fatter tails than doctors and doctors charge a smaller part of patient income.

Welfare inequality. The lack of a uniform quality-adjusted price implies that prices vary along the income distribution. Heterogeneity in consumption patterns implies that people at different points of the income distribution face different price indices (Deaton, 1998). Taking this into account implies that a given increase in income inequality translates into a lower increase in welfare inequality. The assignment mechanism implies that as inequality increases, the rich generalists cannot obtain better health services—in fact they pay more for health services of the same quality. This mechanism limits the welfare increase in inequality. Moretti’s (2013) work on real wage inequality across cities can be viewed as proposing a similar assignment mechanism causing high earners to locate in high-cost cities.⁹

To assess this formally, we use as a consumption-based measure of welfare the homogeneous good consumption $eq(x)$ which, when combined with a fixed level of health quality (namely z_c) gives the same utility to the generalist as what she gets in the market. That is we define $eq(x)$ through $u(z_c, eq(x)) = u(z(x), c(x))$. We then obtain:

Remark 1. For x large enough, the welfare measure eq is Pareto-distributed with shape parameter $\alpha_{eq} \equiv \frac{\alpha_x}{1 + \frac{\alpha_x \beta_z}{\alpha_z (1 - \beta_z)}}$ so that $\frac{d \ln \alpha_{eq}}{d \ln \alpha_x} = \frac{1}{1 + \frac{\alpha_x \beta_z}{\alpha_z (1 - \beta_z)}}$, implying that an increase in inequality for generalists’ income translates into a less than proportional increase in their welfare inequality. The mitigation is stronger when health services matter more (high β_z) or when doctors’ abilities are more unequal (low α_z).

Ssee Appendix A.3 for the proof.

2.1.3 Extensions for empirical testing

In the following sub-sections, we will test the importance of each key assumption in our basic result. We first establish the necessity of the “non-divisibility” assumption by introducing brewers who produce a divisible good, beer, and show that income inequality of brewers is independent of that of generalists.

⁹Diamond’s (2016) critique argues that the amenities of expensive cities are more valuable to the high earners who choose to live there, so we should not fully adjust incomes for these high costs when calculating welfare. In our context, this critique would apply if high-income generalists had stronger preferences for high-quality doctors than low-income generalists.

Second, we show that the predictions of our model are unchanged if we allow mobility across occupations such that high-earning doctors can work as high-earning generalists.

Third, in preparation for our empirical analysis, we introduce a multi-region model. Naturally, without trade or migration between regions top income inequality among doctors must be determined by local generalist income inequality. We show that this remains true even if we allow doctors to move across regions. If we instead allow for the cross-region trade of medical services, income inequality for doctors will be the same for all regions. This distinction between “local” services that cannot be traded across regions and tradable “non-local” services will be important for the empirical section, which is driven by local variation in general income inequality.

Finally, we show robustness of our core result to other model tweaks. Our results hold if the ability distributions are only asymptotically Pareto distributed, and if doctors consume medical services themselves. We use a more general utility function and show that the spillover effect survives, although the prediction of a spillover elasticity of 1 from Proposition 1 does not generalize.

2.2 The Role of the Assortative Matching Mechanism

To highlight the specificity of our mechanism, we add “brewers” to the system. Potential brewers can produce a divisible good, beer. They differ in their ability such that a brewer of ability y can produce quantity y of (quality-adjusted) beer. Their ability distribution is Pareto with shape α_y ; that is a brewer has ability $Y > y$ with probability

$$P(Y > y) = \left(\frac{y_{min}}{y} \right)^{\alpha_y},$$

and α_y is kept constant. If potential brewers do not produce beer they produce x_{min} units of the homogeneous good. We modify the utility function such that $u(z, c, y) = z^{\beta_z} c^{1-\beta_z-\beta_y} y^{\beta_y}$. The first order condition for beer consumption together with a market clearing equation determine the price of beer. As beer is divisible, the beer price p must be taken as given by each producer and brewers’ incomes will simply be given by py . As a result, the income of active beer producers is Pareto distributed with a shape parameter α_y . A change in inequality among generalists can only affect active producers proportionately.¹⁰ Moreover since beer is divisible, the distribution of the “real” income

¹⁰As showed in Appendix A.4, a decrease in α_x increases p for parameters where all potential brewers are actively producing beer. If the extensive margin of brewers is operative the (mean-preserving) in-

inequality is unaffected by the presence of beer and the difference between nominal and real income is only driven by the presence of doctors: Remark 1 still applies and α_{eq} does not change. Consequently, divisibility is essential for spillovers through consumption.

2.3 Occupational Mobility

Above we assumed that a potential doctor working as a generalist makes the minimum amount possible as a generalist: x_{min} . In reality it is quite plausible that those succeeding as doctors would have succeeded in other occupations as well (Kirkeboen et al., 2016). To capture this, we now switch to the opposite extreme and assume that there is perfect correlation between abilities as a doctor and as a generalist. We keep the model as before, except we assume that there is a mass 1 of agents who decide whether they want to be doctors or generalists. We rank agents in descending order of ability and use i to denote their rank, so that the most skilled agent has rank 0 and the least skilled has rank 1. For two agents i and i' with $i < i'$, i will be better both as a generalist and as a doctor than i' . We assume that both ability distributions are Pareto with parameters (x_{min}, α_x) for generalist and (z_{min}, α_z) for doctors. An agent i can choose between becoming a generalist earning $x(i)$ or being a doctor providing health services of quality $z(i)$ and earning $w(z(i))$. Those working as doctors also need the services of doctors. We assume that $\lambda > 1$ to ensure that everyone can get health services. By definition of the rank we have that the counter-cumulative distribution functions (1 minus the CDFs) for x and z obey:

$$\bar{G}_x(x(i)) = \bar{G}_z(z(i)) = i.$$

In equilibrium, it is always the case that below a certain rank, some individuals will choose to be doctors. In addition under parameter conditions detailed below, some individuals will also choose to be generalists (details in Appendix A.7). That is, for i low enough, agents must be indifferent between becoming a doctor or a generalist: $w(z(i)) = x(i)$, which directly implies that, for z high enough, the wage function must satisfy:

$$w(z) = \bar{G}_x^{-1}(\bar{G}_z(z)).$$

Since both ability distributions are Pareto, this can be written as:

crease in income inequality will lower x_{min} and encourage a supply increase of brewers. As a consequence the effect on beer prices, p , from a decrease in α_x is ambiguous.

$$w(z) = x_{\min} \left(\frac{z}{z_{\min}} \right)^{\frac{\alpha_z}{\alpha_x}}. \quad (11)$$

Doctor wages grow in proportion to what they could earn as a generalist.

Let $\mu(z) \in (0, 1)$ denotes the share of individuals able to provide health services of quality z who are doctors. For z sufficiently high that individuals of rank $\bar{G}_z(z)$ and below and their patient work both as generalist and doctors, market clearing implies

$$\left(\frac{x_{\min}}{m(z)} \right)^{\alpha_x} = \int_z^\infty \lambda \mu(\zeta) g_z(\zeta) d\zeta, \quad (12)$$

where $m(z)$ denotes the income (earned either as a generalist or a doctor) of the patient of a doctor of quality z .

The first order condition on health care consumption (3) still applies, and together with (11) and (12) it implies that:

$$\int_z^\infty \mu(\zeta) \alpha_z \zeta^{-\alpha_z - 1} d\zeta = \lambda^{\alpha_x - 1} z^{-\alpha_z} \left(\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1 - \beta_z} \right)^{-\alpha_x}.$$

Differentiating with respect to z , we find that μ is a constant: $\mu = \lambda^{\alpha_x - 1} \left(\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1 - \beta_z} \right)^{-\alpha_x}$. Intuitively, with a constant μ , doctors' wages grow proportionately with patients' incomes, in line with the Cobb-Douglas assumption. Note that since we assumed that $\mu < 1$, this situation is only possible as long as $\lambda^{\alpha_x - 1} \left(\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1 - \beta_z} \right)^{-\alpha_x} < 1$.

Therefore, if $\lambda^{\alpha_x - 1} \left(\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1 - \beta_z} \right)^{-\alpha_x} < 1$, we have that $P_{doc}(W_d > w_d) = P(Z > w^{-1}(w_d))$ for w_d high enough so that the observed distribution for doctor wages is Pareto with a shape parameter α_x : Proposition 1 still applies (in fact the distribution is now exactly Pareto above a threshold). We solve for the full model in Appendix A.7.¹¹ Further, if the distributions of x and z are only asymptotically Pareto, then our results remain true asymptotically, so that Proposition 1 applies.

Note that in terms of observed top income inequality the model where agents can switch and the one where they cannot are observationally equivalent: doctors' top income inequality perfectly traces that of the generalists. This is so because even when doctors are not allowed to shift across occupations, the relative reward to the very best doctors adjusts correspondingly with the shift for generalists.

¹¹If $\lambda^{\alpha_x - 1} \left(\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1 - \beta_z} \right)^{-\alpha_x} > 1$, then all individuals above a certain ability threshold choose to be doctors while all those below it choose to be generalists. This seems counterfactual.

Supply versus demand side effects. In the model just presented doctors and generalists interact both through a demand effect—generalists are the clients of doctors—and a supply effect—doctors can choose to become generalists. Since the wage level is directly determined by doctors’ outside option (according to (11)), one may think that the mechanism which leads to spillovers in income inequality is very different compared to the demand-side mechanism of the baseline model. This is, however, not the case. In Appendix A.8 we split the role of generalists into two: patients, who only serve the role of consumers of doctor services and an “outside option” which only serves the role of providing doctors with an alternative occupation to providing medical services. We show that when the utility function is given by (1), the income inequality of doctors is entirely driven by that of their patients and is independent of changes in the income inequality for the outside option. Consequently, the driving force is still the demand side.

2.4 Mobility and Open Economy

So far we assumed a closed economy. Since our empirical analysis will rely on local variation in income inequality, we next consider an economy with more than one region. We analyze a case in which medical services can be traded between regions and a case in which doctors can move across regions.

2.4.1 Tradable health care

Consider the baseline model of section 2.1. We now assume that there are several regions, $s = 1, \dots, S$ and we allow some patients (a positive share of generalists in all regions) to purchase their medical services across regions. The distribution of potential doctors’ ability is the same in all regions (and so is the parameter λ). The other parameters, and in particular the Pareto shape parameter of generalists’ income α_x^s is allowed to differ across regions. The cost of health care services must be the same everywhere; otherwise, the generalists who can travel would go to the region with the cheapest health care. Since top talented potential doctors work as doctors (instead of being generalists with income x_{\min}^s), they must all earn the same wage. In all regions, the income distribution of patients is asymptotically Pareto with parameter $\min_s \alpha_x^s$, because at the very top, overall income inequality follows the income inequality of the most unequal region. (Section 2.5.1 elaborates on this logic in more depth.) As a result doctors’

income is asymptotically Pareto with shape parameter $\min_s \alpha_x^s$ in all regions. In other words, income inequality for generalists in the *most unequal* region spills over to doctors in *all* regions.

Empirically, whether the service provided is “local” (non-tradable) or “non-local” (tradable) will depend on the occupations of interest. We will use the results of this section and the previous ones to guide our empirical analysis.

2.4.2 Doctors moving

We return again to the baseline model of section 2.1, but we now assume that there are 2 regions, A and B , and that doctors can move across regions.¹² But medical services are again non-tradable and patients cannot move.¹³ The two regions are identical except for the ability distribution of generalists, which is Pareto in both but with possibly different means and shape parameters.¹⁴ Without loss of generality, we assume that $\alpha_x^A < \alpha_x^B$; that is region A is more unequal than region B .

With no trade in goods between the two regions, we can normalize the price of the homogeneous good to 1 in both. As doctors only consume the homogeneous good, doctors’ nominal wages must be equalized in the two regions. As a result the price of health care of quality z must be the same in both regions. From the first order condition on health care consumption, this implies that the matching function is the same: doctors of quality z provide health care to generalists of income $m(z)$ in both regions. Moreover, the least able potential doctor who decides to become a doctor must have the same ability z_c in both regions.

We define by $\varphi(z)$ the *net* share of doctors initially in region B with ability *at least* z who decide to move to region A . Then labor market clearing in region A implies that, for $z \geq z_c$,

$$\left(x_{\min}^A/m(z)\right)^{\alpha_x^A} = \lambda\mu_d(1 + \varphi(z))(z_{\min}/z)^{\alpha_z}. \quad (13)$$

There are initially $\mu_d(z_{\min}/z)^{\alpha_z}$ doctors with ability at least z in each region and by definition, a share $\varphi(z)$ of those move from region B to region A . Since each doctor can provide services to λ patients, after doctors have relocated the total supply over a quality z in region A is given by the right-hand side of (13). Total demand corresponds

¹²The results generalize to more than 2 regions.

¹³When doctors are mobile and medical services tradable, the geographic location of agents is undetermined in general, and we would need a full spatial equilibrium model to generate empirical predictions.

¹⁴Our results directly generalize to a case where the two regions do not have the same mass of potential doctors and generalists.

to region A patients with an income higher than $m(z)$, of which there are $P(X > m(z))$. The same equation, replacing $\varphi(z)$ by $-\varphi(z)$, holds in region B :

$$(x_{\min}^B/m(z))^{\alpha_x^B} = \lambda\mu_d(1 - \varphi(z))(z_{\min}/z)^{\alpha_z}. \quad (14)$$

Since the two regions are of equal size, total demand for health services must be the same and on net, no doctors move: $\varphi(z_c) = 0$. On the other hand, most rich patients are in region A (as $\alpha_x^A > \alpha_x^B$). As doctors' incomes increase with the incomes of their patients, nearly all of the most talented doctors will eventually locate in region A : $\lim_{z \rightarrow \infty} \varphi(z) = 1$. We therefore obtain that, in region A , the distribution of doctors' ability after relocation is asymptotically Pareto. So, as in the baseline model, doctors' incomes will be asymptotically Pareto distributed with a shape parameter equal to α_x^A .

In region B , doctors of a given quality level earn the same as in region A . That is, the incomes of doctors initially in region B are still Pareto distributed with coefficient α_x^A . However, after the move, the share of doctors that stay in region B decreases with their quality. Using (13) and (14), we get that $1 - \varphi(z) \propto z^{\alpha_z(1 - \alpha_x^B/\alpha_x^A)}$. Therefore, the *ex post* talent distribution of doctors in region B is still Pareto but now with a coefficient $\alpha'_z = \alpha_z \alpha_x^B / \alpha_x^A$. As in the baseline model, the distribution of income for doctors who stay must be asymptotically Pareto with a shape Parameter α_x^B .¹⁵ We obtain:

Proposition 2. *Once doctors have relocated, the income distribution of doctors in region A is asymptotically Pareto with coefficient α_x^A , and the income distribution of doctors in region B is asymptotically Pareto with coefficient α_x^B .*

Formal proof is in Appendix A.9.

Consequently, whether doctors can move or not does not alter the *observable* local income distribution, although it does matter considerably for the *unobservable* local *ability* distribution. Consequently, for our empirical analysis we need not take a stand on whether doctors are mobile. We cannot empirically distinguish between the free-mobility and no-migration cases using data on income inequality.

¹⁵To see that there is no contradiction, note that the baseline model predicts that the income of individual z , $w(z) \propto z^{\frac{\alpha'_z}{\alpha_x^B}}$ but $\frac{\alpha'_z}{\alpha_x^B} = \frac{\alpha_z}{\alpha_x^A}$, so we also have $w(z) \propto z^{\frac{\alpha_z}{\alpha_x^A}}$ and doctors do indeed earn the same in both regions.

2.5 Utility Function and Ability Distribution

2.5.1 Doctors consume medical services and ability distribution is only Pareto distributed in the tail

We now alter the model so there is a mass 1 of agents, of which a fraction μ_d are potential doctors. The technology for health services is the same as before (and we now assume that $\lambda > 1/\mu_d$). Agents not working as doctors produce a composite good which we take as the numeraire. Unlike in the baseline model, all agents have the same utility function (1).

The equilibrium results in a wage distribution. We assume that this distribution and also the distributions of skills for potential doctors are asymptotically Pareto. Therefore we can write

$$P_x(X > x) = \bar{G}_x(\bar{x}) \bar{G}_{x,\bar{x}}(x),$$

where $\bar{G}_{x,\bar{x}}(x)$ is the conditional counter-cumulative distribution above \bar{x} and $\bar{G}_x(\bar{x})$ is the unconditional counter-cumulative distribution, and for \bar{x} large enough we have

$$\bar{G}_x(x, \bar{x}) \approx (\bar{x}/x)^{\alpha_x},$$

with $\alpha_x > 1$. The same holds for doctors' talents z (moreover, potential doctors can work as generalists with the lowest productivity x_{\min} as an alternative).

As before, solving for the consumer problem leads to the differential equation (3). Furthermore since health care services are not divisible, the equilibrium also features assortative matching and we still denote the matching function $m(z)$. Market clearing at every z can still be written as (4). The least able potential doctor who actually works as a doctor will have ability $z_c = \bar{G}_z^{-1}(1/\lambda\mu_d)$. Therefore z_c is independent of α_x . As a result, (4) implies that $m(z)$ is defined by $m(z) = \bar{G}_x^{-1}(\bar{G}_{z,z_c}(z))$.

For z above some threshold, \bar{z} , both doctors' talents and incomes are approximately Pareto distributed, which allows us to rewrite the previous equation as:

$$\bar{G}_x(m(\bar{z})) (m(\bar{z}) / (m(z)))^{\alpha_x} \approx \bar{G}_{\bar{z},z_c}(\bar{z}) (\bar{z}/z)^{\alpha_z},$$

which gives

$$m(z) \approx B z^{\frac{\alpha_z}{\alpha_x}} \text{ with } B = m(\bar{z}) \left(\frac{\bar{G}_x(m(\bar{z}))}{\bar{G}_{\bar{z},z_c}(\bar{z}) \bar{z}^{\alpha_z}} \right)^{\frac{1}{\alpha_x}}.$$

Plugging this in (3) we can rewrite the differential equation as:

$$w'(z)z + \frac{\beta_z}{1 - \beta_z}w(z) \approx \frac{\beta_z}{1 - \beta_z}\lambda Bz^{\frac{\alpha_z}{\alpha_x}}.$$

Therefore for z large enough, we must have (see Appendix A.5 for a derivation):

$$w(z) \approx \frac{\beta_z \alpha_x}{\alpha_z (1 - \beta_z) + \beta_z \alpha_x} \lambda B z^{\frac{\alpha_z}{\alpha_x}}. \quad (15)$$

From this we get (as above) that for \bar{w}_d large enough, doctors' income is distributed according to

$$P(W_d > w_d | w_d > \bar{w}_d) \approx \left(\frac{\bar{w}_d}{w_d} \right)^{\alpha_x}. \quad (16)$$

That is, doctors' income follows a Pareto distribution with shape parameter α_x . Proposition 1 still applies: a decrease in α_x will directly translate into an increase in top income inequality among doctors.

2.5.2 The role of the Cobb-Douglas utility function

We keep the same model as just introduced, but we replace the utility function of equation (1) with:

$$u(z, c) = \left(\beta_z z^{\frac{\varepsilon-1}{\varepsilon}} + \beta_c c^{\frac{\varepsilon-1}{\varepsilon}} \right)^{\frac{\varepsilon}{\varepsilon-1}}, \quad (17)$$

with $\varepsilon \neq 1$. As before, the first order conditions gives the differential equation:

$$\partial u / \partial z = \omega'(z) \partial u / \partial c. \quad (18)$$

Since CES exhibits positive cross-partial derivatives, we know that the equilibrium features positive assortative matching. Therefore, with income and ability asymptotically Pareto, the matching function still obeys (5). Using (17), combining (18) and (5), and using that $w(z) = \lambda \omega(z)$, we find that for high levels of z the wage function obeys a differential equation given by

$$w'(z) \approx \lambda^{\frac{\varepsilon-1}{\varepsilon}} \frac{\beta_z}{\beta_c} z^{-\frac{1}{\varepsilon}} \left(\lambda B z^{\frac{\alpha_z}{\alpha_x}} - w(z) \right)^{\frac{1}{\varepsilon}}. \quad (19)$$

We solve this differential equation in Appendix A.6 and we prove:

Proposition 3. *i) Assume that $\varepsilon > 1$. Then for $\alpha_x \geq \alpha_z$, wages of doctors are asymp-*

totically Pareto distributed with exponential parameter $\alpha_w = \alpha_x$. For $\alpha_x < \alpha_z$, wages of doctors are asymptotically Pareto distributed with $\alpha_w = \frac{\alpha_z}{\left(\frac{\alpha_z}{\alpha_x} - 1\right)^{\frac{1}{\varepsilon} + 1}}$.

ii) Assume that $\varepsilon < 1$. Then for $\alpha_x > \frac{\alpha_z}{1 - \varepsilon}$, wages of doctors are bounded. For $\alpha_x = \frac{\alpha_z}{1 - \varepsilon}$, wages of doctors are asymptotically exponentially distributed. For $\alpha_z < \alpha_x < \frac{\alpha_z}{1 - \varepsilon}$, they are asymptotically distributed with $\alpha_w = \frac{\alpha_z}{\left(\frac{\alpha_z}{\alpha_x} - 1\right)^{\frac{1}{\varepsilon} + 1}}$. For $\alpha_x \leq \alpha_z$, they are asymptotically Pareto distributed with $\alpha_w = \alpha_x$.

Therefore, when doctors' income distribution is Pareto, we still obtain that a reduction in α_x leads to a reduction in α_w (that is an increase in general top income inequality increases top income inequality among doctors), although the elasticity may now be lower than 1. (It cannot be asymptotically above 1, since high-paying generalists would then spend more than their income on medical services.) Further, a decrease in α_x also reduces the size of the parameter space for which doctors' wage distribution is bounded (a situation where top income inequality for doctors is very low).

To intuitively understand the results of Proposition 3, consider first the case where $\alpha_z > \alpha_x$. That is, the ability distribution of generalists has a fatter tail than that of doctors, implying a shortage of doctors at the top. This must mean a convex pricing schedule for medical services. If $\varepsilon > 1$, health services and the homogeneous good are substitutes, so the expenditure share on health services declines with income. As a result, $w(z)$ cannot grow as fast as the income of the generalist who buys the services of doctor z , namely $m(z)$, which grows as $z^{\frac{\alpha_z}{\alpha_x}}$. This implies less income inequality among the top doctors than among the top generalists (a higher Pareto exponential parameter). On the other hand, if $\varepsilon < 1$ then richer generalists are forced to spend an increasing amount—eventually all their resources—on health services. So $m(z)$ and $w(z)$ grow at the same rate, and doctors' income is Pareto distributed with coefficient α_x . The reverse holds when doctors are relatively abundant at the top (i.e. when $\alpha_z < \alpha_x$), except that with $\varepsilon < 1$, doctors' income can even be bounded.

2.6 Empirical predictions

To summarize, our model makes the following predictions:

1. An increase in general inequality will lead to an increase in inequality for doctors if they service the general population directly and their services are non-divisible.
2. This is true regardless of whether doctors can move across regions, and regardless

of whether doctors’ ability is positively correlated with the income they would receive in alternative occupations.

3. If patients can easily travel, doctors’ income in each region does not depend on local income inequality.

3 Empirical Strategy and Data

3.1 Empirical strategy

We are centrally interested in the causal effect of general top income inequality in a region s on the top income inequality of a particular subgroup i in region s . Since our data are top-censored, we make the distributional assumption that the right tail of the income distribution is Pareto distributed: $P(X > x) = (x/x_{min})^{-\alpha}$ above some cut-off x_{min} and use $1/\alpha$ as a our measure of income inequality. Specially, for such a distribution the relative income of somebody at the 99th percentile relative to somebody at the 95th percentile is $5^{1/\alpha}$ and the Gini coefficient is $(2\alpha - 1)^{-1}$. Guvenen, Karahan, Ozkan, and Song (2015) and Jones and Kim (2014) also employ $1/\alpha$ as a measure of income inequality.

Using this, the regression of interest is:

$$\log\left(\frac{1}{\alpha_{o,t,s}}\right) = \gamma_s + \gamma_t + \beta \log\left(\frac{1}{\alpha_{-o,t,s}}\right) + X_{t,s}\delta + \epsilon_{o,t,s}, \quad (20)$$

where $\frac{1}{\alpha_{o,t,s}}$ is top income inequality for occupation i at time t for geographical area s and $\frac{1}{\alpha_{-o,t,s}}$ is the corresponding value for the general population in s except for i . Let γ_s be a dummy for the geographical area, γ_t a time dummy, and $X_{t,s}$ a vector of controls, including the area’s population and average income. We are centrally interested in β which measures the elasticity of top income inequality for our occupation of interest with respect to the general income inequality. We will focus on labor market areas, which are aggregations of commuting zones (Dorn, 2009) and can generally be driven through in a matter of a few hours, e.g. Los Angeles or New York. Central results carry through if we instead use commuting zones or states as the unit of analysis.

We consider a number of occupations, but are limited by the fact that our analysis requires a relatively high number of observations among a high-earning population to measure inequality. We split these occupations into two groups: First, we focus on

occupations whose output is non-divisible and who primarily operate in local markets: physicians, dentists and real estate agents. Admittedly, some patients do travel for special medical treatment. To the extent that this creates an integrated market, this would create a downward bias in our estimate of the spillover effects. In the extreme, section 2.4 showed that full tradability eliminates local spillovers.

We contrast these with other occupations who have also seen increases in income inequality, but that do not satisfy these conditions: college professors, secretaries, and financial managers. Although professors’ output may be non-divisible, they do not operate in a local market—at least not in the right tail of the distribution. Although secretaries operate in local markets, they do not service the general population directly. Finally, consider financial managers. According to the Standard Occupational Classification scheme, financial managers “plan, direct, or coordinate accounting, investing, banking, insurance, securities, and other financial activities of a branch, office, or department of an establishment.” We think of them in two categories: those who manage financial matters for corporations and who are unlikely to be affected by higher local income inequality, and those that manage the financial means of private individuals who are likely to operate in more integrated markets. In either case we expect to see no local spillover effect. Figure 2 shows the increase in the 99th/90th percentile income ratios for the selected occupations. It demonstrates that the increase in within-occupation top income inequality is a trend outside of the very top of the income distribution as well.

We will estimate regression (20) by using the publicly available Decennial Census and American Community Survey to estimate both $1/\alpha_{o,t,s}$ and $1/\alpha_{-o,t,s}$. This allows us to examine the time period 1980 to 2014 and consider a relatively broad set of occupations. Due to possible concerns about endogeneity we will use an instrumental variable approach, using a “shift-share” instrument (following Bartik, 1991) based on the occupational distribution across geographical areas in 1980. We first show summary statistics in Section 3.3. We then perform our regression analysis on the income data in Section 4. Appendix B graphically shows how well the data fit the Pareto distribution.

3.2 Income data

Our central data set is a combination of the Decennial Census for 1980, 1990 and 2000 and the American Community Survey (ACS) for 2010-2014 (which, combined together,

we refer to as 2014) (Ruggles et al. 2015).¹⁶ We have 5.4 million observations in 1980, growing to 7.4 million observations in 2014, with positive wage income. We use 2010-2014 as opposed to the perhaps more natural 2008-2012 to avoid the immediate aftermath of the Great Recession, which had large impact on top income. Data from farther back are a substantially smaller sample so we exclude them from the analysis. We use the 1990 census occupational classification from IPUMS, which consistently assigns occupations throughout the 1980-2014 period. The publicly available income data are censored, generally at around the 99.5th percentile of the overall income distribution, which complicates our estimation of the parameter of the Pareto distribution.¹⁷ In particular, suppose \tilde{X} follows a Pareto distribution $P(\tilde{X} > \tilde{x}) = (\tilde{x}/x_{min})^{-\alpha}$, but the observed wage is $x = \min\{\tilde{x}, \bar{x}\}$ for some censoring point, \bar{x} . Then we can write the maximum likelihood function as $\mathcal{L}(\alpha) = (\prod_{i \in \mathcal{N}_{unc}} \alpha (x_{min}/x_i)^\alpha x_i^{-1}) (x_{min}/\bar{x})^{\alpha N_{cen}}$, where \mathcal{N}_{unc} is the set of uncensored observations and N_{cen} is the number of censored observations. Armour, Burkhauser and Larrimore (2014) use the same methodology on the Current Population Survey (March supplement) to show that trends in income inequality match those found by Kopczuk, Saez and Song (2010) using uncensored social security data. The resulting maximum likelihood estimate is

$$\frac{1}{\hat{\alpha}} = \frac{\sum_{i \in \mathcal{N}_{unc}} \log(x_i/x_{min}) + N_{cen} \log(\bar{x}/x_{min})}{N_{unc}}, \quad (21)$$

where N_{unc} is number of uncensored observations. Note, that even without the assumption of a Pareto distribution, equation (21) is a measure of income inequality: It is the average log-difference from the minimum possible observation for the uncensored observations, plus the product of the relative number of censored observations times the log distance from the censoring point to the minimum. This will be our measure of income inequality throughout.

Since we need a reasonable number of observations in order to estimate local inequality, we restrict ourselves to the biggest 253 labor market areas — those with at least 8 observations of physicians in 1980 — for a total of 1,012 observations. Furthermore, to account for the fact that the key variables in our regressions are themselves estimated, we

¹⁶The Decennial Censuses are each 5 per cent of the population, whereas the ACS each are 1 per cent. Combining the years 2010-2014 creates an ‘artificial’ sample of 5 per cent for 2014. The IPUMS inflates all numbers to 2014 using the consumer price index.

¹⁷Specially, the censoring takes place at \$75,000 for 1980, \$140,000 for 1990, \$175,000 for 2000 and at the 99.5th percentile at the state level for each individual year 2010-2014. The information provided about the censored variables varies from year to year.

calculate confidence intervals using bootstrapping. We bootstrap from the microdata, re-estimating $1/\hat{\alpha}$ in each replication. We stratify the data at the occupational-labor market area-year level and use 300 replications.¹⁸

3.3 Summary statistics

We will be focusing on the combined pre-tax wage and salary income throughout.¹⁹ Table 1 below shows the mean, median, 90th, 95th and 98th percentiles among those with positive wage income for each year. All values are in 2014 dollars.²⁰ As discussed in the introduction, the ratios of the 98th to 95th percentiles, and the 95th percentile to the median, have increased during the period. The table also shows the estimate of $1/\alpha$ on the top 10 per cent of observations with positive wage income for each year. We present the 98th/95th and 95th/90th ratios implied by the Pareto distribution with the estimated $\hat{\alpha}$ in parentheses. There is a high level of agreement between the predicted and the actual ratios, consistent with a good fit to the Pareto distribution.

Although the censoring point is sufficiently high to allow standard measures of top income inequality to be calculated for most occupations, the high average income of some occupations leads to a larger share being censored. Although the censoring has little impact on the overall distribution, slightly more than 26 per cent of Physicians with positive income are censored in 2000. This implies that we cannot calculate measures of income inequality using high percentiles. But we can still calculate $1/\alpha$ using the assumption of a Pareto distribution. Table 2 shows the result using the top 65 per cent of the uncensored observations.²¹ Consistent with Figure 1, $1/\alpha$ has increased for most occupations in the top during this period. Table B.1 in the Appendix shows the

¹⁸That is for each draw we resample person-observations, recalculate the α 's and reestimate the regressions. For computational reasons we do restrict attention to the top end of the redistribution, i.e. we only resample from top 10 per cent of the income distribution in a given labor market area to calculate general income inequality.

¹⁹The census includes other measures of income, in particular business income, which could be relevant for some occupations. Unfortunately, since wage income and business income are censored separately, estimating a joint distribution for the two would be substantially more complicated. We are in process of getting access to the full uncensored data which would allow us to use total income.

²⁰We use the 98th percentile as the censoring doesn't allow for the calculations of 99th for all years.

²¹Throughout the paper we follow the following rule of thumb when calculating occupation, year, labor market specific measures of income inequality: If there are very few censored observations —say for secretaries— we use the top 10 per cent of the distribution. For occupations that are heavily censored — physicians and dentists — we move the cut-off until we have around twice as many uncensored observations as censored for all labor market areas we use. For Physicians that is the top 65 per cent, for dentists it is top 50 per cent and for Real Estate agents it is top 20 per cent.

calculated measures of income inequality (using the top 10 per cent of the population) for a number of other occupations, along with the fraction of observations with positive income that are censored. The table shows the same general trend, but with some notable exceptions. In particular there has been little upward trend in top income inequality for truck drivers, sales people, and computer software developers, but substantial increases for financial managers and chief executives. Table B.2 shows which occupations were in the top 1, 5 and 10% for 1980 and 2014. What is particularly noteworthy is that Physicians are increasingly important in the high end of the income distribution and that this importance has grown from 1980 to 2014. In fact, Physicians are the most common (Census) occupation in the top 1% in 2014.

Each observation in the data is associated with a particular geographical area (“county groups” for 1980 and “Public Use Microdata Areas” from 1990 onward). These are statistical areas created to ensure confidentiality and, according to Dorn (2009), have little economic meaning. Alternatively, one could use states, but some local economies, say greater New York City or Washington D.C. span several states. At the same time, some states are too large to meaningfully capture a local economy and others are too small to have sufficient number of observations. Dorn instead uses commuting zones, of which there are 741. We take a similar approach but use labor market areas. Both commuting zones and labor market areas are defined based on the commuting patterns between counties (Tolbert and Sizer, 1996). But whereas commuting zones are unrestricted in size, labor market areas aggregate commuting zones to ensure a population of at least 100,000. Given that our estimation strategy relies on a relatively high number of observations of a particular occupation, labor market areas are a more natural choice. Table 3 shows the size distribution of number of total observations with positive wage income and physicians with positive wage income across labor market areas.

To assess the fit of the Pareto distribution at the labor market area, year, occupation level, we use the fact that a Pareto distribution implies a linear relationship between value and frequency. Figure 3 shows this relationship for the biggest labor market area for both all occupations (Los Angeles) and physicians specifically (New York). The line shows the predicted number of observations in each bin and the orange dots give the actual number of observations in each bin, both plotted according to the left-side y axis. The right-side y axis gives the corresponding values for the censored values, scaled to ensure that the predicted censored values are on the same line as the uncensored

predicted values.²²

Figure 3.a below uses the biggest labor market area (Los Angeles) for the year 2000 and bins the income interval between that of the 90th percentile and the censoring point of 175,000 into 20 evenly sized bins and plots the (linear) predicted number of observations from the associated Pareto distribution with the observed number of observations in each bin (the choice of bins in the figure does not influence any estimation results). The figure further shows the actual and predicted number of censored observations on the right hand side scaled to fit a linear line. The fit for the general population is very close to a straight line and therefore a Pareto distribution. We perform an analogous analysis for the physicians (where New York City is the biggest labor market), although with the lower number of observations overall, and the much higher number of censored observations, we use the top 65 per cent of the positive uncensored observations.²³ The fewer observations implies a fit that is less tight, but there are no systematic deviations from the straight line. Figures B.1 and B.2 in the Appendix give equivalent figures for the 20 biggest labor markets in the United States.

3.4 Instrument

One might worry about endogeneity when estimating equation (20). In particular, even controlling for labor market area and year fixed effects, a positive correlation between general income inequality and income inequality for a specific occupation might reflect deregulation, changes in the tax system or common local economic trends and not reflect a causal effect from general income inequality to inequality for the occupation of interest. To address this issue, we use a Bartik (1991)-style instrument. We define:

$$\log I_{-o,t,s} = \log \left[\sum_{\kappa \in K_{-o}} \omega_{\kappa,1980,s} (1/\alpha_{\kappa,t}) \right], \text{ for } t = 1980, 1990, 2000, 2014.$$

²²Formally, we use the fact that for a dataset with N observations on wages drawn from a Pareto distribution $P(X > x) = (x/x_{min})^{-\alpha}$ with a corresponding pdf of $f(x) = \alpha x^{-(\alpha+1)} x_{min}^{\alpha}$, the expected number of observations that have wage income in the interval $[x' - \frac{\Delta}{2}, x' + \frac{\Delta}{2}]$ is $N_{x'} = N \int_{x' - \frac{\Delta}{2}}^{x' + \frac{\Delta}{2}} f(x) dx \simeq N \Delta \alpha x^{-(\alpha+1)} x_{min}^{\alpha}$, giving a negative linear relationship between $\log N_{x'}$ and $\log x$. The predicted number of censored observations is $P(X > \bar{x}) = (\bar{x}/x_{min})^{-\alpha}$ to which we (arbitrarily) assign the value $\bar{x} + \Delta/2$ and scale to fit on the same predicted line.

²³Though we carry out the main analysis using the top 65% of observations, Table C.1 in the Appendix shows that the parameter estimate is relatively insensitive to the choice of cut-off.

K_{-o} is the set of the 20 most important occupations in the top 5 per cent of the income distribution nationwide in 1980 (excluding occupation o). $\omega_{\kappa,1980,s}$ is the share of individuals in occupation κ among individuals in an occupation belonging to K_{-o} in 1980 in LMA s . Goldsmith-Pinkham et al. (2017) show that this amounts to using 1980 industry shares as instruments, with a weighting matrix determined by changes in nationwide occupation-specific inequality. The instrument has strong predictive power: the correlation between $\log I_{-o,t,s}$ and $\log(1/\alpha_{-o,t,s})$ for Physicians is between 0.39 and 0.55 for each year (depending on the number of LMAs considered). It is practically the same for all occupations as each occupation represents a relatively small share of total top income holders.²⁴ In other words, in the IV regression we only exploit the changes in labor market income inequality that arises from the occupational distribution in 1980 combined with the nationwide trends in occupational inequality. Furthermore, by using nationwide trends, our instrument is more likely to capture the effects of globalization, technological change or deregulation, which affect local inequality but are exogenous to the LMA, which is in line with a decrease in α_x in our theoretical model.

4 Empirical Analysis

4.1 Testing the model for occupations with positive predicted spillovers

Having estimated the Pareto distributions described above we next estimate equation (20). We start out by conducting the analysis for physicians. Table 4 presents summary statistics for the regressors of interest.

The core result is shown in Table 5.²⁵ The first column shows an OLS regression of physicians' income inequality on general income inequality including year and LMA fixed effects. We find an elasticity of around 1/3. This estimate remains unchanged in column (2), where we include controls for labor market population and the average wage income among those with positive wage income. Neither control has a significant impact on physicians' income inequality. Column (3) shows the first stage of the instrumental variable regression using the instrument as constructed in (21). The instrument has a

²⁴The qualitative conclusions of our analysis remain unchanged by using a different number of top occupations than 20, although the point estimate of β is somewhat sensitive.

²⁵One can show that the inverse of the variance of the MLE estimator of equation 21 is proportional to the number of uncensored observations and we correspondingly weigh the equation by the number of uncensored observations of physicians.

strong predictive power and, along with the time trends, accounts for 82 percent of the variation in the variation for general income inequality (R^2 is computed excluding the LMA fixed effects). The F -statistic for the first stage is 40 for our preferred specification and higher than 20 for all regressions in this paper (as recommended by Montiel Olea and Pflueger, 2013).

The fourth and fifth columns give the main IV results, which show point estimates of 1.18 or 1.42 for the coefficient of interest, depending on controls. This is strongly significantly different from 0, and not significantly different from the value of 1 predicted by the simplest model in section 2.1. With our measure of general top income inequality increasing by 27 percent since 1980, and top income inequality for physicians increasing by 31 percent (Tables 1 and 2), an elasticity of 1 suggests that a large share of the rise of income inequality among doctors can be explained by the general increase in income inequality, although the exact fraction is measured with uncertainty. Moreover, note that neither the controls nor the year fixed effects are significant in the IV regressions, which is consistent with our mechanism explaining most of the changes in income inequality for doctors.

In Appendix Table C.2, we show that the results are robust to dramatic changes in sample size. We change the size cut-off between the top 100 LMAs and all LMAs where we are able to estimate physician inequality. The parameter estimate, β , remains significant and between 0.8 and 1.7.

We next analyze two other occupations that are much less regulated and even more local: dentistry and real estate. We perform an analogous examination of dentists in Table 6 and reach broadly similar conclusions. Again we focus on labor market areas with at least 8 observations in 1980; this severely reduces the number of labor market areas from 253 to 40. Yet we see a pattern broadly similar to that of physicians, albeit with less precision (the OLS coefficients have p-values of 0.12). Both OLS and IV point estimates are around twice as high for dentists as for physicians. Though this might reflect the fact that dentistry is more local and prices are less regulated, the point estimates are not significantly distinct and we cannot rule out a difference purely due to sampling error. With a spillover elasticity of 2.8 and a rise in income inequality for dentists that has mirrored that of the general population we substantially “over-explain” the rise in income inequality for dentists, though with this few observations there is substantial imprecision in the estimate.²⁶

²⁶We also perform the analogous analysis on nurses for whom top inequality has grown as well (See Figure 2) though it is less clear that our model would apply to this occupation. Whereas the OLS

Finally, we use an occupation outside the medical industry: real estate agents. The fee structure in real estate is often proportional to housing prices (Miceli et al., 2007) and the increase in the spread of housing prices is consistent with the increase in income inequality (Määttänen and Terviö, 2014). Real estate is a difficult business to scale up, as each house still needs to be shown individually and each transaction negotiated separately. Consequently, one would expect to see spillover effects from general income inequality to real estate agents. Table 7 shows that this is indeed the case. Though the OLS estimates are somewhat lower than for the physicians, the IV estimates are very close. Income inequality for Real Estate agents has increased from 0.45 to 0.69, an increase of 50%. With general income inequality increasing by around 27% the IV estimate suggests that more than half the increase in agents’ income inequality can be attributed to the general increase in income inequality.

4.2 Testing the model when spillovers are not predicted

Whereas our theory predicts local spillover effects from general income inequality to the income inequality for occupations such as physicians, dentists and real agents, it predicts no such spillovers for other occupations. We perform analogous regressions for financial managers, who, as argued above, do not fit the conditions required for local spillovers. Table 8 shows that this is the case. Though the OLS estimate is positive, the IV estimate is close to zero (and in fact the point estimate is now negative). This also shows that spurious correlation between general inequality and occupational inequality at the local level is likely but that our instrument can address this concern.

Finally, we perform the analysis for two other occupations with substantial increases in top income inequality from Figure 2 but where our model predicts no spillovers: College Professors and Secretaries.²⁷ Tables 10 and 11 show these results. The top 10% of earners among university professors operate in a national market, and secretaries are not hired by private citizens. For these occupations we find no effect.

Intriguingly, income inequality for secretaries correlates strongly with income inequality for “Chief Executives and Public Administrators” in an OLS regression, shown

estimates are similar to physicians, the IV estimate is lower and estimated with more imprecision and we cannot clearly establish spillovers as a cause for the increase in income inequality for nurses. See Table 9 for details.

²⁷There is a data break in the IPUMS data: For 1980 to 1990 Post-Secondary teachers (those teaching at higher level than high-school) are partly categorized by subject of instruction (code 113-154). From 2000 onward they are not. We collapse all codes 113-154 into 154 for 1980 to 1990.

in Appendix Table C.3.) Though we cannot establish causality, this is consistent with a theory analogous to ours, in which CEOs compete for the most skilled secretaries. But at slightly above 10%, the implied elasticity is substantially lower than for other occupations considered here.

5 Physician Pricing and Networks

To test the mechanism for spillovers that the model proposes, we delve into the details of how physician price-setting works in practice in the United States. We exploit data on privately negotiated physician payments and the structure of insurers' networks to directly test the mechanism proposed in our model.

5.1 Institutions

For multiple reasons, the medical industry in the United States is not perfectly described by the perfectly flexible price-setting model of section 2.1. The government plays a substantial role through Medicare and Medicaid, the insurance sector has an important role as an intermediary, there is substantial asymmetric information between patients and doctors, and patients are often willing to travel to seek medical attention. But these features need not substantially impact our analysis. Although the government sets administrative prices for those whose care it pays for directly, providers' negotiations with private insurers generally lead to higher prices (Clemens and Gottlieb, 2017). Even in the presence of asymmetric information, patients often have clear beliefs about who the "best" local doctor in a specific field is (whether or not these beliefs relate to medical skill or health outcomes). And although patients occasionally travel for care, a patient in Dallas is vastly more likely to seek medical care in Dallas than Boston. Furthermore, our empirical strategy more heavily weights large metropolitan areas, which are more likely to have a full portfolio of medical specialties implying less need to travel. To the extent that the medical industry is best described by a national market, the model suggests that this will simply reduce our estimated spillovers.

Despite these complications, the structure of the health insurance industry may embody enough flexibility to incorporate the economic pressures implied by our model. Clemens, Gottlieb and Molnár (2017) show that insurers and physicians frequently negotiate reimbursements as fixed markups over Medicare. They find that, if Medicare

sets a reimbursement rate of r_j^M for treatment j , private insurer i 's reimbursement to physician group g for that treatment is generally determined by

$$r_{i,g,j} = \varphi_{i,g} r_j^M.$$

Following their logic, we will use the markups $\varphi_{i,g}$ as a summary measure of the prices charged by physician group g for treating insurer i 's patients. Again following Clemens, Gottlieb and Molnár (2017), we estimate these markups with a regression of the form

$$\ln r_{i,g,j} = \phi_{i,g} + \ln r_j^M + \varepsilon_{i,g,j} \tag{22}$$

on insurance claims data—data that record insurers' payments to provider groups for specific treatments. In equation (22), $\phi_{i,g}$ is an insurer-physician fixed effect which we interpret as the log of the group's markup over Medicare rates. Clemens, Gottlieb and Molnár (2017) show that this regression matches realized physician payments extremely well, and that the levels of these markups reflect economic pressures such as physician market power.

Nevertheless, in practice these markups are not determined in a completely decentralized market, with each physician setting a price to implement the perfectly assortative matching that our model contemplates. Instead, patients purchase insurance and insurers group beneficiaries into different plans, distinguished largely by the breadth of their networks. That is, an expensive Gold plan may have a large network encompassing most physicians in a region, while a cheaper Silver plan may pay physicians lower reimbursements and have a smaller network (Polsky et al., 2016).

This network structure provides a mechanism to mediate the heterogeneous consumer preferences that inequality generates. A consumer with a high willingness to pay for physician quality would have to buy an expensive plan, which pays high physician reimbursements. Because of these high reimbursements, many physicians agree to join the plan's network and treat that plan's customers.²⁸ A consumer with a lower willingness to pay can buy a cheaper plan, which saves money by paying physicians less—and, as a result, fewer physicians join that network. So the lower-willingness-to-pay patient ends up with less choice of physicians.

²⁸The insurer enforces the network by providing different levels of coverage when patients see in-network and out-of-network providers. Patients who visit an out-of-network physician normally have to pay more, or even all of the cost, out of pocket. In contrast, those who see the in-network physicians that have agreed to accept the network's reimbursement rates generally incur little or no out-of-pocket cost.

We will use data on physician networks to test whether this mechanism is more pronounced in areas with more inequality. Based on the logic of our model, income inequality should predict more variability in network breadth and physicians’ network participation. This provides an institutionally-informed mechanism to transmit income inequality into physician price inequality.

5.2 Data and empirical approach

Insurance claims data We measure inequality in physician payments using three source of insurance claims data. The first two sources are the same ones used in Clemens, Gottlieb and Molnár (2017), and are fully described there: Blue Cross/Blue Shield of Texas (BCBS-TX) and the Colorado All-Payer Claims Data (APCD-CO). As a third source, we add All-Payer Claims Data from New Hampshire (APCD-NH). All three datasets have a similar structure: they provide details on patient visits for physician care, and indicate the service provided and the identity of the physician providing treatment (as an actual name or in encrypted form). Crucially, they indicate the amount the physician was paid for each service, the insurer providing coverage, and whether the physician is in-network. Relying on the institutional details described above, we focus on in-network payments.

Depending on the details of the patient’s insurance contract, and whether the patient has reached an annual deductible or out-of-pocket maximum, the patient or the insurer may have to pay the physician’s fee for a particular treatment. But regardless of who is liable, the amount that the physician expects to receive is governed by the rate negotiated between the physician and the insurer. The three databases all provide information on this negotiated amount, known in the industry as the “allowed charge.” They indicate the treatment that the fee covers using a 5-digit code established by the Healthcare Common Procedure Coding System (HCPCS).

This provides the information necessary to estimate equation (22). In the BCBS-TX data, there is only one insurer so we simply employ physician group fixed effects. In the other datasets, we decompose $\phi_{i,g}$ additively into physician and insurer fixed effects. We estimate (22) on each dataset, which provides us with a distribution of log prices. We use these to compute local inequality measures of physician prices. We compute the same $\log(1/\alpha)$ measure we have used throughout the paper, as well as ratios of the 90th to 50th and 75th to 50th percentiles of these markups.

The BCBS-TX data encompass years from 2008-2013, so we will primarily use them

as a panel and take differences between the local inequality measures in 2008 and 2013. We regress this short difference on the change in the Bartik instrument from 2000 to 2014, the closest pair of years available for that measure. When using this instrument we continue to run the analysis at the LMA level.

Since we have shorter panels for the APCD datasets, we run a three-state analysis as a pure cross-section. We amalgamate all of the years of data to form one cross-section, but add richness by computing inequality measures at the finer commuting zone level. We then regress physician price inequality on local income inequality excluding physicians.

Insurance network data To study inequality in physician networks, we use data collected by the Narrow Networks Project (NNP) at the University of Pennsylvania (Polsky and Weiner, 2015; Zhu et al., 2017). This dataset lists the physicians participating in each insurance network for the health insurance exchange plans established under the Affordable Care Act. It reports the physician’s identifier, location, and plan participation. We combine it with data on the total number of physicians in each county from the Area Resource File, a standard reference produced by the Department of Health and Human Services.

We construct two primary measures from these datasets. First we consider the share of physicians participating in any exchange plan at all. Since the exchange plans tend to pay lower reimbursements than standard private insurance plans, this is effectively an inverse measure of physicians accepting only high-paying patients. In other words, it measures the uniformity of the health insurance market in a region.

Our second summary is a more direct inequality measure. For each insurance network in a region, we count the number of physicians participating in that network according to the NNP data. We then compute the standard deviation of this measure across networks, which we then normalize by the mean to have a coefficient of variation (CV). This directly measures variability in the size of a network, which provides a mechanism for transmitting heterogeneity in patients’ willingness-to-pay into heterogeneity in physician reimbursements.

These measures are only available for a short time horizon: three years, in which the ACA exchanges were just starting and were constantly in flux. So we only use the most recent year’s data and treat them as a cross-section. We regress the network inequality measures on our standard inequality measure for non-physicians in an LMA,

$\log(1/\alpha_{-o,t,s})$.

We standardize all of the network measures, as well as our inequality measure, so regression coefficients can be easily interpreted in terms of standard deviations.

5.3 Results

We report the physician pricing results in Table 12. Columns 1-4 show reduced form regressions of changes in pricing inequality across Texas LMAs from 2008-2013 against changes in the Bartik instrument. The coefficient of 1.6 in column 1 is statistically indistinguishable from the baseline result for physician income in Table 5. Based on this result, a one-standard-deviation increase in the instrument (0.016) would lead to a one-quarter-standard-deviation increase in pricing inequality growth (0.026/0.112). Column 2 adds controls for the specialty composition in an area, which leads both the coefficient and standard error to approximately double. Results are similar when we use pricing ratios in columns 3 and 4, although column 4 loses statistical significance.

Columns 5 and 6 turn to cross-sectional regressions on a larger sample: three states, and with data at the commuting zone level. These coefficients are not directly comparable to the earlier columns, as the dependent and independent variables are now levels rather than prices. Furthermore, the independent variable is now the realized local income inequality rather than the instrument. These coefficients imply somewhat smaller standardized results: a one-standard deviation increase in inequality (0.115) is associated with one-sixth of a standard deviation higher prices (0.086/0.517) according to column 5. Nevertheless the association remains strongly positive. More unequal areas, and areas with growing predicted inequality, experience more inequality in physician reimbursements.

Table 13 presents the network inequality results. The coefficient of 0.72 in column 1 means that variability of network size is 0.74 standard deviations higher in an LMA with 1-standard-deviation higher inequality. Column 2 adds controls for the mix of specialties in an area, and the coefficient falls slightly to 0.62. Both coefficients are statistically significant. Columns 3 and 4 turn to the extensive margin of physician participation in ACA exchange networks. Here, we find that 1-standard-deviation higher inequality is associated with a 0.42-standard-deviation fall in ACA network participation, or 0.23 standard deviations after adding controls. Figure 4 shows these results graphically, using a binned scatterplot.

6 Conclusion

In this paper, we established that an increase in income inequality in one occupation can spill over through consumption to other occupations, such as physicians, dentists and real estate agents, that provide non-divisible services directly to customers. We show that changes in general income inequality at the level of the local labor market area do indeed spill over into these occupations. We distinguish this mechanism by considering other occupations that have seen rises in top income inequality, but that either do not fit our assumptions or operate in a national labor market. Financial managers and college professors experience no spillover effects. This alligns clearly with the predictions of our theory of consumption-driven spillovers. Data on the specific operation of physician markets provides further support for our mechanism.

The magnitude of the key results suggests that this effect may explain most of the increase in income inequality for occupations such as doctors, dentists and real estate agents. As a result, the increase in top income inequality across most occupations observed in the last 40 years may not require a common explanation. Increases in inequality for, say, financial managers or CEOs because of deregulation or globalization may have spilled over to other high-earning occupations, causing a broader increase in top income inequality.

This analysis has been purely positive, but clearly has normative implications, which we plan on exploring in future work. In particular, our analysis could be relevant to the study of top income taxation (see for instance, Scheuer and Werning, 2015).

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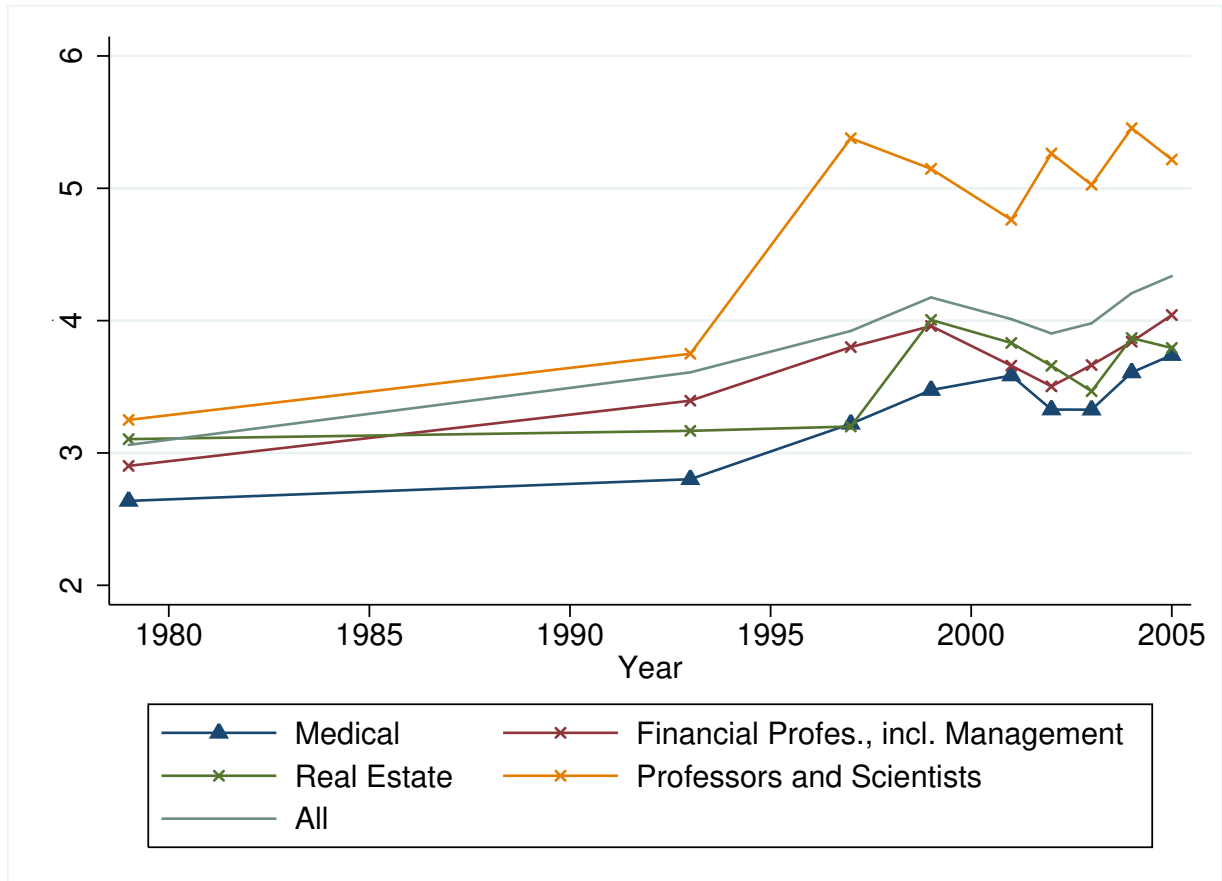
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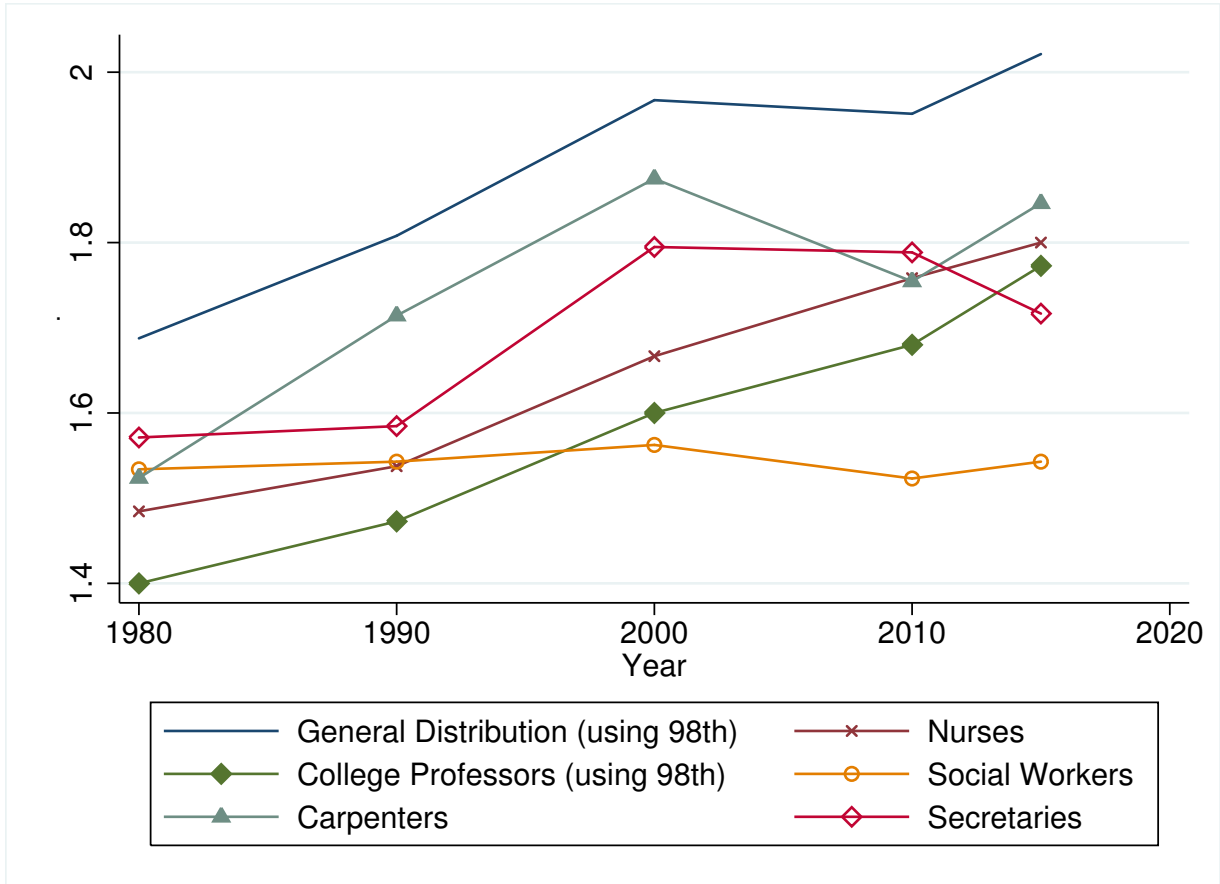
Zhu, J. M., Zhang, Y., and Polsky, D. (2017). Networks in ACA marketplaces are narrower for mental health care than for primary care. *Health Affairs*, 36(9):1624–1631.

Figure 1: Relative Income: Top 0.1% to Top 1% for Selected Occupations



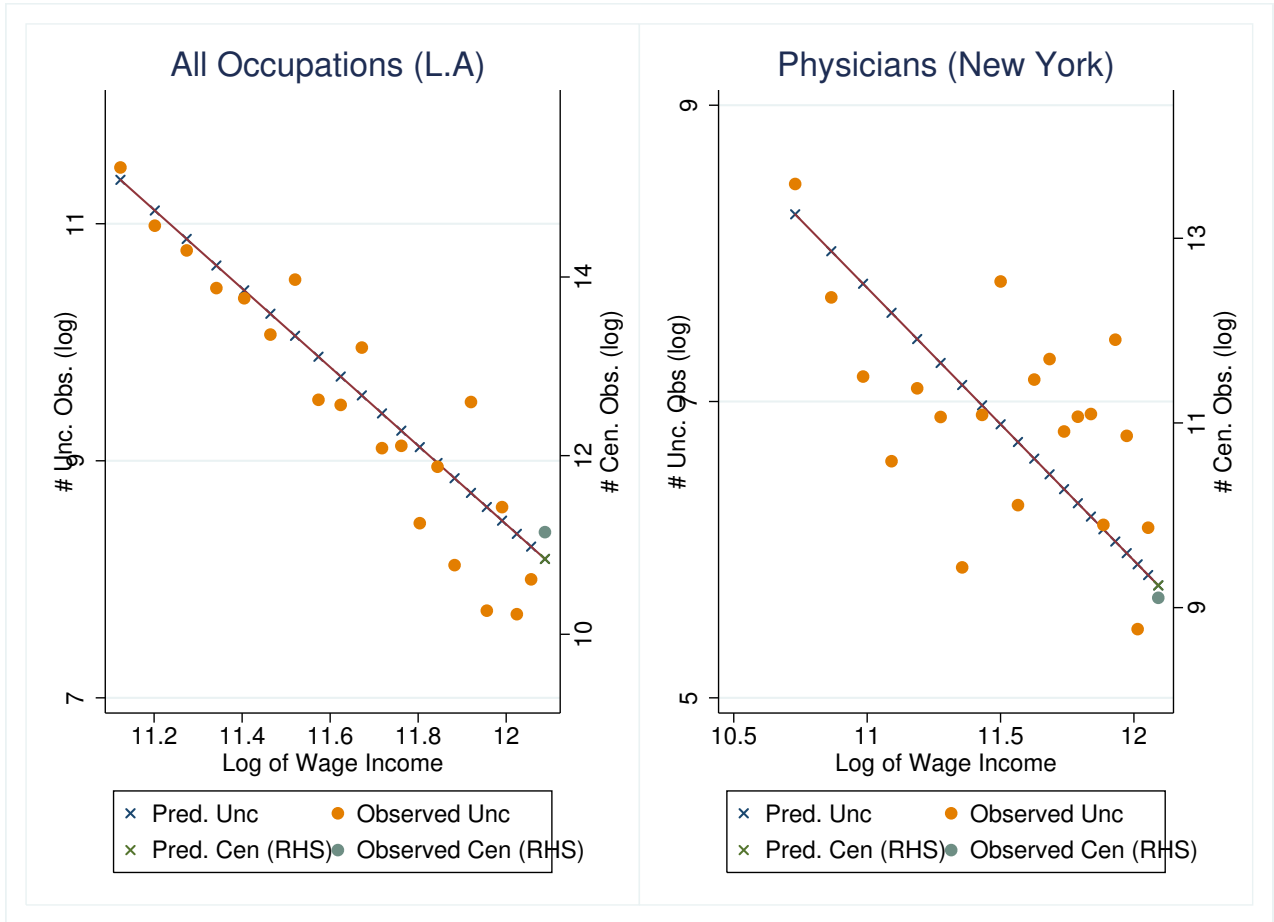
Notes: This figure shows the ratio of mean earnings among those in the top 0.1% of the income distribution relative to those in the top 1%, for selected occupations. “All” refers to the full income distribution, not just the occupations shown here. Source: Bakija, Cole and Heim (2012)

Figure 2: Relative Income: 99th to 90th Percentile for Selected Occupations



Notes: This figure shows the ratio of the 99th percentile of the income distribution to the 90th percentile, for selected occupations. The sample consists of employed workers with positive wage income. Censoring prevents the calculation of the 99th percentile for the general distribution as well as for college professors, so we show the 98th percentile instead for those samples. Source: Authors' calculations using Decennial Census and American Community Survey data from IPUMS (Ruggles et al. 2015).

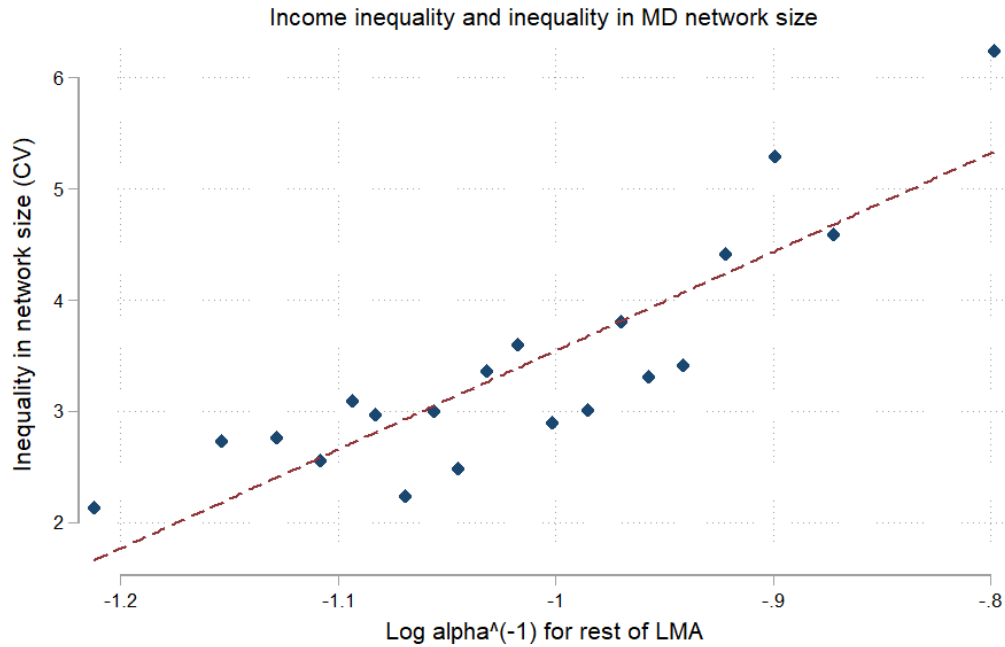
Figure 3: Fit of the Pareto Distribution



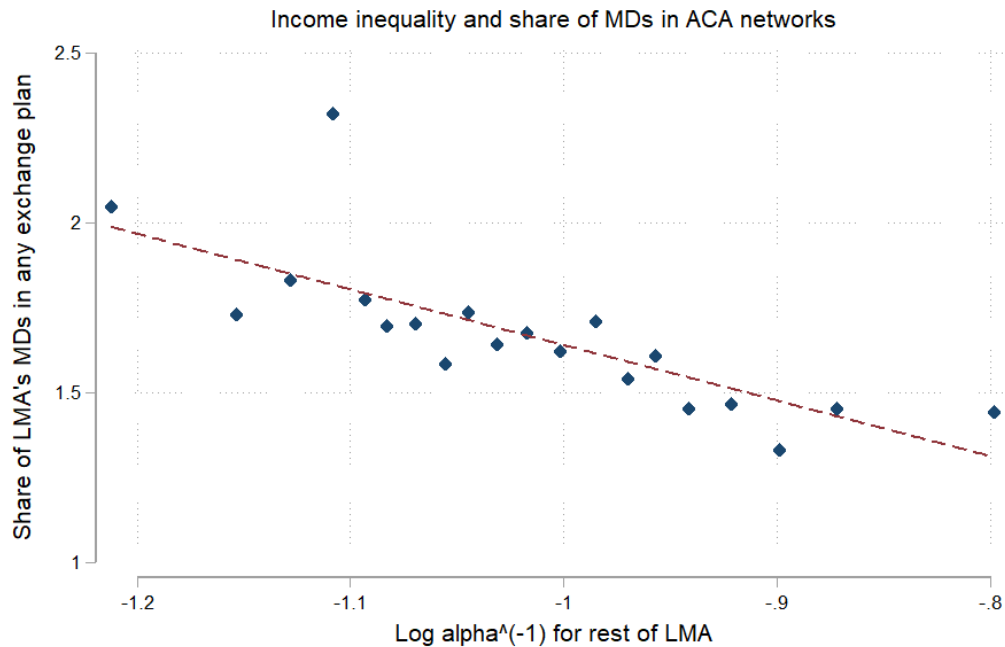
Notes: This figure shows the quality of fit of the empirical income distribution to the Pareto distribution for two samples in data from 2000. The left panel shows the full sample in Los Angeles, and the right panel shows physicians in New York. The line shows the predicted distribution if it were entirely Pareto, and the dots show the empirical sample sizes. The color changes for the highest income value shown, which includes the observations where income is censored. Source: Authors' calculations using Decennial Census data from IPUMS (Ruggles et al. 2015).

Figure 4: Inequality and Physician Network Structure

Panel A



Panel B



Notes: Panel A shows the relationship between income inequality in an LMA (excluding physicians) and inequality in local ACA network size. Panel B shows the relationship between income inequality and the share of local physicians participating in any ACA network plan. In both cases, we group the data into twenty sized bins based on local income inequality. Sources: Income inequality is constructed from 2014 Census data provided by IPUMS (Ruggles et al. 2015). Physician network measures are based on authors' calculations from the University of Pennsylvania's Narrow Networks Project (Polsky and Weiner, 2015) and the Area Resource File.

Table 1: Wage income 1980-2014 for general population

Year	Median	percentile			p95/p90 (predicted)	p98/p95 (predicted)	1/ α
		p90	p95	p98			
1980	25.9	68.1	85.2	114.9	1.25 (1.26)	1.35 (1.35)	0.33
1990	29.0	76.1	96.8	137.6	1.27 (1.30)	1.42 (1.42)	0.38
2000	33.0	83.9	111.4	165.0	1.33 (1.32)	1.48 (1.44)	0.40
2014	30.5	90.0	120.0	177.9	1.33 (1.34)	1.48 (1.47)	0.42

Notes: Real wage income for observations with positive income (1000s of 2014 dollars using CPI). p95/p90 is the relative income of top 5 and top 10 per cent (predicted values in parentheses). Top-censoring prevents the calculation of 99th percentile wages. $1/\alpha$ are exponential parameters of a Pareto distribution and are calculated by MLE

Table 2: Wage income 1980-2014 for Physicians

Year	median	1/ α	p95/p90 (pred)
1980	100.6	1.29	2.44
1990	126.8	1.49	2.80
2000	137.5	1.58	2.99
2014	160.9	1.70	3.26

Notes: Real Wage Income (1000s of 2014 dollars using CPI). α is exponential parameter of a Pareto distribution and is calculated through MLE using the top 65 per cent of non-censored positive income

Table 4: Summary Table For Regression Variables

Variable	Obs.	Mean	Std. dev.	Min	Median	Max
$\log(1/\alpha(doc))$	1,012	0.28	0.50	-2.17	0.33	2.34
$\log(1/\alpha(else))$	1,012	-1.11	0.14	-1.60	-1.10	-0.70
$\log(I)$	1,012	-1.15	0.10	-1.52	-1.13	-0.91

Notes: For labor market areas where top 65 per cent uncensored physicians includes at least 8 observations

Table 3: Number of observations across labor market areas

Year	Mean	All			Physicians			
		25th	Median	75th	Mean	25th	Median	75th
1980	39584	17500	24560	37269	70	17	28	54
1990	43772	21207	29392	43286	89	22	37	70
2000	47541	21829	31786	46297	105	26	42	87
2014	51495	22789	34212	49489	138	29	56	115

Notes: Number of observations for labor market areas (all and physicians)

Table 5: Regression Table for Physicians

	(1)	(2)	(3)	(4)	(5)
	OLS	OLS	1st Stage	2SLS	2SLS
	$\log(1/\alpha(o))$	$\log(1/\alpha(o))$	$\log(1/\alpha(-o))$	$\log(1/\alpha(o))$	$\log(1/\alpha(o))$
$\log(1/\alpha(-o))$	0.32*** [0.10, 0.55]	0.28** [0.04, 0.50]		1.18*** [0.67, 1.65]	1.42*** [0.74, 1.99]
$\log(I)$			1.76*** [1.72, 1.79]		
\logpop		-0.05 [-0.18, 0.05]	-0.06*** [-0.06,-0.05]		0.12 [-0.04, 0.26]
$\log(\text{income})$		0.13 [-0.05, 0.29]	0.04*** [0.04, 0.05]		0.12 [-0.06, 0.28]
LMA FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
R2 (ex. LMA FE)	0.47	0.48	0.89	.	.
Observations	184	184	184	184	184

Notes: This table shows the full set of baseline results for income inequality spillovers to physicians from other occupations in the top 20 among top income earners. Columns 1 and 2 show the OLS relationship between local income inequality among physicians and among the rest of the population in a given labor market area. Column 3 shows the first stage relationship between the Bartik instrument, using inequality among the top 20 non-physician occupations, and local non-physician income inequality. Columns 4 and 5 show the second stage of the two-stage least squares estimate using that same instrument. In column 3 only, the dependent variable is income inequality among non-physicians in an LMA. In the remaining columns, the dependent variable is income inequality among physicians in an LMA, while non-physician inequality is the main right-hand-side variable. In all cases, inequality is measured as the log inverse of the Pareto parameter ($\log(1/\alpha)$). The occupation of interest is denoted with o . Standard errors are bootstrapped using 300 draws, stratified at the occupation-year-labor market level. Square brackets show the 95 percent confidence interval based on this bootstrapping. Statistical significance is denoted by * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The income control is log average wage income for those with positive income.

Table 6: Regression Table for Dentists

	(1)	(2)	(3)	(4)	(5)
	OLS	OLS	1st Stage	2SLS	2SLS
	$\log(1/\alpha(o))$	$\log(1/\alpha(o))$	$\log(1/\alpha(-o))$	$\log(1/\alpha(o))$	$\log(1/\alpha(o))$
$\log(1/\alpha(-o))$	0.60 [-0.12, 1.21]	0.54 [-0.16, 1.20]		2.17* [-0.11, 3.79]	2.77* [-0.19, 5.20]
Instrument			1.66*** [1.20, 2.04]		
Log of Population		-0.04 [-0.35, 0.35]	-0.12*** [-0.14,-0.08]		0.30 [-0.21, 0.83]
Log of Income		0.42 [-0.15, 0.85]	0.03 [-0.01, 0.08]		0.38 [-0.20, 0.87]
LMA FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
$R^2(\text{ex. LMA FE})$	0.19	0.20	0.87	.	.
Observations	160	160	160	160	160

Notes: This table shows the full set of baseline results for income inequality spillovers to dentists from other occupations in the top 20 among top income earners. Columns 1 and 2 show the OLS relationship between local income inequality among dentists and among the rest of the population in a given labor market area. Column 3 shows the first stage relationship between the Bartik instrument, using inequality among the top 20 non-dentist occupations, and local non-dentist income inequality. Columns 4 and 5 show the second stage of the two-stage least squares estimate using that same instrument. In column 3 only, the dependent variable is income inequality among non-dentists in an LMA. In the remaining columns, the dependent variable is income inequality among dentists in an LMA, while non-dentist inequality is the main right-hand-side variable. In all cases, inequality is measured as the log inverse of the Pareto parameter ($\log(1/\alpha)$). The occupation of interest is denoted with o . Standard errors are bootstrapped using 300 draws, stratified at the occupation-year-labor market level. Square brackets show the 95 percent confidence interval based on this bootstrapping. Statistical significance is denoted by * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The income control is log average wage income for those with positive income.

Table 7: IV Regressions for Real Estate Agents (top 20 per cent)

	(1)	(2)	(3)	(4)	(5)
	OLS	OLS	1st Stage	2SLS	2SLS
	$\log(1/\alpha(o))$	$\log(1/\alpha(o))$	$\log(1/\alpha(-o))$	$\log(1/\alpha(o))$	$\log(1/\alpha(o))$
$\log(1/\alpha(-o))$	0.17*	0.17*		1.02**	1.32**
	[-0.03, 0.30]	[-0.03, 0.30]		[0.20, 2.09]	[0.29, 2.56]
Instrument			0.64***		
			[0.51, 0.76]		
Log of Population		0.05*	-0.04***		0.11***
		[-0.00, 0.10]	[-0.05,-0.03]		[0.04, 0.20]
Log of Income		0.23***	0.03**		0.19**
		[0.13, 0.33]	[0.00, 0.05]		[0.08, 0.31]
LMA FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
R^2 (ex. LMA FE)	0.05	0.05	0.82	.	.
Observations	1,448	1,448	1,448	1,448	1,448

Notes: This table shows the full set of baseline results for income inequality spillovers to real estate agents from other occupations in the top 20 among top income earners. Columns 1 and 2 show the OLS relationship between local income inequality among realtors and among the rest of the population in a given labor market area. Column 3 shows the first stage relationship between the Bartik instrument, using inequality among the top 20 non-realtor occupations, and local non-realtor income inequality. Columns 4 and 5 show the second stage of the two-stage least squares estimate using that same instrument. In column 3 only, the dependent variable is income inequality among non-realtors in an LMA. In the remaining columns, the dependent variable is income inequality among realtors in an LMA, while non-realtor inequality is the main right-hand-side variable. In all cases, inequality is measured as the log inverse of the Pareto parameter ($\log(1/\alpha)$). The occupation of interest is denoted with o . Standard errors are bootstrapped using 300 draws, stratified at the occupation-year-labor market level. Square brackets show the 95 percent confidence interval based on this bootstrapping. Statistical significance is denoted by * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The income control is log average wage income for those with positive income.

Table 8: Regression Table for Financial Managers (top 10 per cent)

	(1)	(2)	(3)	(4)	(5)
	OLS	OLS	1st Stage	2SLS	2SLS
	$\log(1/\alpha(o))$	$\log(1/\alpha(o))$	$\log(1/\alpha(-o))$	$\log(1/\alpha(o))$	$\log(1/\alpha(o))$
$\log(1/\alpha(-o))$	0.77*** [0.21, 1.34]	0.63** [0.01, 1.26]		1.29** [0.15, 2.64]	1.02 [-0.81, 2.78]
$\log(I)$			1.27*** [1.20, 1.30]		
\logpop		-0.21 [-0.50, 0.13]	-0.08*** [-0.09,-0.07]		-0.16 [-0.61, 0.23]
$\log(\text{income})$		0.13 [-0.27, 0.56]	0.01 [-0.00, 0.01]		0.14 [-0.26, 0.58]
LMA FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
R2 (ex. LMA FE)	0.46	0.47	0.89	.	.
Observations	184	184	184	184	184

Notes: This table shows the full set of baseline results for income inequality spillovers to financial managers from other occupations in the top 20 among top income earners. Columns 1 and 2 show the OLS relationship between local income inequality among financial managers and among the rest of the population in a given labor market area. Column 3 shows the first stage relationship between the Bartik instrument, using inequality among the top 20 non-financier occupations, and local non-financier income inequality. Columns 4 and 5 show the second stage of the two-stage least squares estimate using that same instrument. In column 3 only, the dependent variable is income inequality among non-financiers in an LMA. In the remaining columns, the dependent variable is income inequality among financial managers in an LMA, while non-financier inequality is the main right-hand-side variable. In all cases, inequality is measured as the log inverse of the Pareto parameter ($\log(1/\alpha)$). The occupation of interest is denoted with o . Standard errors are bootstrapped using 300 draws, stratified at the occupation-year-labor market level. Square brackets show the 95 percent confidence interval based on this bootstrapping. Statistical significance is denoted by * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The income control is log average wage income for those with positive income.

Table 9: IV Regressions for Nurses (top 10 per cent)

	(1)	(2)	(3)	(4)	(5)
	OLS	OLS	1st Stage	2SLS	2SLS
	$\log(1/\alpha(o))$	$\log(1/\alpha(o))$	$\log(1/\alpha(-o))$	$\log(1/\alpha(o))$	$\log(1/\alpha(o))$
$\log(1/\alpha(-o))$	0.37 [-0.14, 0.75]	0.32 [-0.17, 0.77]		0.07 [-1.09, 1.14]	-0.11 [-1.64, 1.46]
$\log(I)$			1.09*** [1.08, 1.12]		
\logpop		-0.07 [-0.27, 0.11]	-0.10*** [-0.10,-0.10]		-0.14 [-0.46, 0.16]
$\log(\text{income})$		0.36*** [0.09, 0.65]	-0.00 [-0.00, 0.00]		0.35** [0.07, 0.64]
LMA FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
R^2 (ex. LMA FE)	0.59	0.61	0.89	.	.
Observations	184	184	184	184	184

Notes: This table shows the full set of baseline results for income inequality spillovers to nurses from other occupations in the top 20 among top income earners. Columns 1 and 2 show the OLS relationship between local income inequality among nurses and among the rest of the population in a given labor market area. Column 3 shows the first stage relationship between the Bartik instrument, using inequality among the top 20 non-nursing occupations, and local non-nursing income inequality. Columns 4 and 5 show the second stage of the two-stage least squares estimate using that same instrument. In column 3 only, the dependent variable is income inequality among non-realtors in an LMA. In the remaining columns, the dependent variable is income inequality among nurses in an LMA, while non-nurse inequality is the main right-hand-side variable. In all cases, inequality is measured as the log inverse of the Pareto parameter ($\log(1/\alpha)$). The occupation of interest is denoted with o . Standard errors are bootstrapped using 300 draws, stratified at the occupation-year-labor market level. Square brackets show the 95 percent confidence interval based on this bootstrapping. Statistical significance is denoted by * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The income control is log average wage income for those with positive income.

Table 10: IV Regressions for College Professors (top 10 per cent)

	(1)	(2)	(3)	(4)	(5)
	OLS	OLS	1st Stage	2SLS	2SLS
	$\log(1/\alpha(o))$	$\log(1/\alpha(o))$	$\log(1/\alpha(-o))$	$\log(1/\alpha(o))$	$\log(1/\alpha(o))$
$\log(1/\alpha(-o))$	0.18 [-0.31, 0.65]	0.34 [-0.13, 0.89]		0.44 [-0.82, 1.53]	0.73 [-0.77, 2.23]
$\log(I)$			1.28*** [1.27, 1.31]		
\logpop		0.22*** [0.04, 0.46]	-0.07*** [-0.07,-0.07]		0.27** [0.01, 0.59]
$\log(\text{income})$		-0.21 [-0.50, 0.16]	0.07*** [0.07, 0.08]		-0.22 [-0.52, 0.15]
LMA FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
R^2 (ex. LMA FE)	0.70	0.71	0.89	.	.
Observations	184	184	184	184	184

Notes: This table shows the full set of baseline results for income inequality spillovers to professors from other occupations in the top 20 among top income earners. Columns 1 and 2 show the OLS relationship between local income inequality among professors and among the rest of the population in a given labor market area. Column 3 shows the first stage relationship between the Bartik instrument, using inequality among the top 20 non-professor occupations, and local non-professor income inequality. Columns 4 and 5 show the second stage of the two-stage least squares estimate using that same instrument. In column 3 only, the dependent variable is income inequality among non-professors in an LMA. In the remaining columns, the dependent variable is income inequality among professors in an LMA, while non-professor inequality is the main right-hand-side variable. In all cases, inequality is measured as the log inverse of the Pareto parameter ($\log(1/\alpha)$). The occupation of interest is denoted with o . Standard errors are bootstrapped using 300 draws, stratified at the occupation-year-labor market level. Square brackets show the 95 percent confidence interval based on this bootstrapping. Statistical significance is denoted by * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The income control is log average wage income for those with positive income.

Table 11: IV Regressions for Secretaries (top 10 per cent)

	(1)	(2)	(3)	(4)	(5)
	OLS	OLS	1st Stage	2SLS	2SLS
	$\log(1/\alpha(o))$	$\log(1/\alpha(o))$	$\log(1/\alpha(-o))$	$\log(1/\alpha(o))$	$\log(1/\alpha(o))$
$\log(1/\alpha(-o))$	0.03 [-0.11, 0.21]	0.06 [-0.09, 0.24]		0.39 [-0.69, 1.38]	0.41 [-0.68, 1.30]
Instrument			0.69*** [0.57, 0.81]		
Log of Population		0.13*** [0.05, 0.22]	-0.06*** [-0.07,-0.04]		0.16*** [0.04, 0.27]
Log of Income		0.16*** [0.03, 0.31]	0.02*** [0.00, 0.04]		0.15** [0.01, 0.31]
LMA FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
R^2 (ex. LMA FE)	0.13	0.14	0.80	.	.
Observations	1,576	1,576	1,576	1,576	1,576

Notes: This table shows the full set of baseline results for income inequality spillovers to secretaries from other occupations in the top 20 among top income earners. Columns 1 and 2 show the OLS relationship between local income inequality among secretaries and among the rest of the population in a given labor market area. Column 3 shows the first stage relationship between the Bartik instrument, using inequality among the top 20 non-secretary occupations, and local non-secretary income inequality. Columns 4 and 5 show the second stage of the two-stage least squares estimate using that same instrument. In column 3 only, the dependent variable is income inequality among non-secretaries in an LMA. In the remaining columns, the dependent variable is income inequality among secretaries in an LMA, while non-secretary inequality is the main right-hand-side variable. In all cases, inequality is measured as the log inverse of the Pareto parameter ($\log(1/\alpha)$). The occupation of interest is denoted with o . Standard errors are bootstrapped using 300 draws, stratified at the occupation-year-labor market level. Square brackets show the 95 percent confidence interval based on this bootstrapping. Statistical significance is denoted by * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The income control is log average wage income for those with positive income.

Table 12: Inequality and Physician Pricing Dispersion

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable (prices):	$\Delta \log \alpha^{-1}$	$\Delta \log \alpha^{-1}$	$\Delta 75/50$ ratio	$\Delta 90/50$ ratio	$\log \alpha_{prices}^{-1}$	$\log \alpha_{prices}^{-1}$
$\Delta \log(I)$	1.641*	3.423*	0.962**	2.723		
	(0.797)	(1.619)	(0.441)	(2.378)		
$\log(\alpha_{incomes}^{-1})$					0.746***	0.516**
					(0.251)	(0.243)
Log population					-0.099***	-0.184***
					(0.016)	(0.029)
Log income						1.190***
						(0.342)
N	20	20	20	20	79	79
R^2	0.06	0.34	0.16	0.02	0.89	0.91
SD of Dep. Var.	0.112	0.112	0.039	0.283	0.517	0.517
SD of Indep. Var.	0.016	0.016	0.016	0.016	0.115	0.115
States	TX	TX	TX	TX	TX,CO,NH	TX,CO,NH
Controls (specialty comp.)	No	Yes	No	No	No	No

Notes: This table shows short-difference and cross-sectional regressions of inequality in physician reimbursements against local income inequality. Columns 1-4 use BCBS-TX pricing data and short differences, while columns 5 and 6 add in Colorado and New Hampshire APCD data and treat the data as a cross-section. The dependent variable in columns 1, 2, 5, and 6 is our standard inequality measure: $\log(\alpha^{-1})$ but calculated using physician markups. Columns 3 and 4 use the 75/50 and 90/50 ratio of physician markups, respectively. The independent variable in columns 1-4 is the change in the value of the Bartik instrument from 2000-2014 (still using as weights the local occupational distribution in 1980), while in columns 5 and 6 we use the 2014 local inequality (excluding physicians). The geographic unit in columns 1-4 is Labor Market Areas in Texas, and in columns 5-6 is Commuting Zones in all three states. Sources: Income inequality is constructed from Census data provided by IPUMS (Ruggles et al. 2015). Physician pricing inequality measures are based on authors' calculations from BCBS-TX, APCD-CO, and APCD-NH data.

Table 13: Inequality and Physician Network Structure

	(1)	(2)	(3)	(4)
Dependent variable (z -scores):	SD network size	SD network size	Network participation	Network participation
$\log(1/\alpha)$ for non-physicians (z -score)	0.724*** (0.110)	0.616*** (0.114)	-0.420*** (0.070)	-0.225*** (0.068)
N	253	253	253	253
R^2	0.25	0.31	0.10	0.36
SD of Dep. Var.	258.47	258.47	0.56	0.56
SD of $\log(1/\alpha)$	0.14	0.14	0.14	0.14
Controls (specialty comp.)	No	Yes	No	Yes

Notes: This table shows cross-sectional regressions of network inequality measures against local income inequality. The dependent variable in columns 1 and 2 is the standard deviation of the number of physicians in local ACA networks. In columns 3 and 4, it is the share of local physicians participating in any local ACA exchange plan network. We standardize both the left- and right-hand-side variables for ease of interpretation. The standard deviations of the original (non-standardized) variables are provided in the table. Sources: Income inequality is constructed from Census data provided by IPUMS (Ruggles et al. 2015). Physician network data are based on authors' calculations from the University of Pennsylvania's Narrow Networks Project (Polsky and Weiner, 2015) and the Area Resource File.

A Appendix: Theory

A.1 Positive assortative matching in equilibrium

Here we show that the equilibrium must feature positive assortative matching between the income of the patient and the skill of the doctor. To do so, we assume that there are 2 individuals 1 and 2 with income $x_1 < x_2$ whose consumption bundles are so that $z_1 > z_2$ and $c_1 < c_2$. For simplicity we write the utility function as a function of health services and the income left for other goods ($x - \omega(z)$).

Note that since consumer 1 chooses a doctor of quality z_1 , it must be the case that:

$$u(z_1, x_1 - \omega(z_1)) \geq u(z_2, x_1 - \omega(z_2)).$$

Further, we have:

$$\begin{aligned} & u(z_1, x_2 - \omega(z_1)) - u(z_2, x_2 - \omega(z_2)) \\ &= u(z_1, x_2 - \omega(z_1)) - u(z_1, x_1 - \omega(z_1)) + u(z_1, x_1 - \omega(z_1)) - u(z_2, x_1 - \omega(z_2)) \\ & \quad + u(z_2, x_1 - \omega(z_2)) - u(z_2, x_2 - \omega(z_2)) \\ &= \int_{x_1 - \omega(z_1)}^{x_2 - \omega(z_1)} \left(\frac{\partial u}{\partial c}(z_1, c) - \frac{\partial u}{\partial c}(z_2, c) \right) + u(z_1, x_1 - \omega(z_1)) - u(z_2, x_1 - \omega(z_2)). \end{aligned}$$

If the utility function has a positive cross-partial (which is the case for a Cobb-Douglas), then the first term is positive as $z_1 > z_2$. Since the second term is also weakly positive, then it must be the case that $u(z_1, x_2 - \omega(z_1)) > u(z_2, x_2 - \omega(z_2))$, in other words, consumer 2 would rather pick a doctor of ability z_1 . Therefore there is a contradiction and it must be the case that $z_1 < z_2$.

A.2 Solving (6)

We look for a specific solution to(6) of the type $w(z) = K_1 z^{\frac{\alpha_z}{\alpha_x}}$. We find that such a K_1 must satisfy

$$K_1 = x_{\min} \frac{\beta_z \alpha_x \lambda}{\alpha_z (1 - \beta_z) + \beta_z \alpha_x} \left(\frac{1}{z_c} \right)^{\frac{\alpha_z}{\alpha_x}}.$$

As the solutions to the differential equation $w'(z)z + \frac{\beta_z}{1-\beta_z}w(z) = 0$ are given by $Kz^{-\frac{\beta_z}{1-\beta_z}}$ for any constant K . We get that all solutions to (6) take the form:

$$w(z) = \frac{x_{\min}\beta_z\alpha_x\lambda}{\alpha_z(1-\beta_z) + \beta_z\alpha_x} \left(\frac{z}{z_c}\right)^{\frac{\alpha_z}{\alpha_x}} + Kz^{-\frac{\beta_z}{1-\beta_z}}.$$

We then obtain (7) by using that $w(z_c) = x_{\min}$ which fixes

$$K = x_{\min}z_c^{\frac{\beta_z}{1-\beta_z}} \frac{\alpha_z(1-\beta_z) + \beta_z\alpha_x(1-\lambda)}{\alpha_z(1-\beta_z) + \beta_z\alpha_x}.$$

A.3 Proof of Remark 1

Using (1), (2), (5) and (10), we get that the utility of a generalist with income x is given by

$$\begin{aligned} u(x) &= (x - h(x))^{1-\beta_z} (m^{-1}(x))^{\beta_z} \\ &= \left(\frac{\alpha_z(1-\beta_z)}{\alpha_z(1-\beta_z) + \beta_z\alpha_x} x - \frac{1}{\lambda} \frac{\alpha_z(1-\beta_z) + \beta_z\alpha_x(1-\lambda)}{\alpha_z(1-\beta_z) + \beta_z\alpha_x} x_{\min} \left(\frac{x}{x_{\min}} \right)^{-\frac{\alpha_x}{\alpha_z} \frac{\beta_z}{1-\beta_z}} \right) \\ &\quad \cdot \left(z_c \left(\frac{x}{x_{\min}} \right)^{\frac{\alpha_x}{\alpha_z}} \right)^{\beta_z}. \end{aligned}$$

Therefore $eq(x)$ obeys

$$\begin{aligned} eq(x) &= \left(\frac{\alpha_z(1-\beta_z)}{\alpha_z(1-\beta_z) + \beta_z\alpha_x} x - \frac{1}{\lambda} \frac{\alpha_z(1-\beta_z) + \beta_z\alpha_x(1-\lambda)}{\alpha_z(1-\beta_z) + \beta_z\alpha_x} x_{\min} \left(\frac{x}{x_{\min}} \right)^{-\frac{\alpha_x}{\alpha_z} \frac{\beta_z}{1-\beta_z}} \right) \\ &\quad \cdot \left(\frac{x}{x_{\min}} \right)^{\frac{\alpha_x}{\alpha_z} \frac{\beta_z}{1-\beta_z}}, \end{aligned}$$

which implies that for x large enough

$$eq(x) \approx \frac{\alpha_z(1-\beta_z)x_{\min}^{-\frac{\alpha_x}{\alpha_z} \frac{\beta_z}{1-\beta_z}}}{\alpha_z(1-\beta_z) + \beta_z\alpha_x} x^{1 + \frac{\alpha_x}{\alpha_z} \frac{\beta_z}{1-\beta_z}}.$$

Then the distribution of real income obeys $\Pr(EQ > e) = \Pr(X > eq^{-1}(e))$, so that for e large enough, we obtain:

$$\Pr(EQ > e) \approx \left(\frac{x_{\min} \alpha_z (1 - \beta_z)}{\alpha_z (1 - \beta_z) + \beta_z \alpha_x e} \right)^{\frac{\alpha_x}{1 + \frac{\alpha_x \beta_z}{\alpha_z (1 - \beta_z)}}}.$$

Therefore asymptotically, real income is distributed in a Pareto way with a shape parameter $\alpha_{eq} \equiv \frac{\alpha_x}{1 + \frac{\alpha_x \beta_z}{\alpha_z (1 - \beta_z)}}$. Moreover we obtain: $\frac{d \ln \alpha_{eq}}{d \ln \alpha_x} = \frac{1}{1 + \frac{\alpha_x \beta_z}{\alpha_z (1 - \beta_z)}}$.

A.4 Brewers' case

Taking first order conditions with respect to c and y , we obtain that expenditures on beers and on the homogeneous good are related by

$$py = \frac{\beta_y}{1 - \beta_y - \beta_z} c. \quad (23)$$

The first order condition with respect to the quality of the health services consumed and the homogeneous good similarly imply

$$\omega'(z) z = \frac{\beta_z}{1 - \beta_y - \beta_z} c. \quad (24)$$

Together with the budget constraint equation

$$\omega(z) + py + c = x,$$

(23) and (24) give (3) so that all results concerning $w(z)$ including (10) still apply, and

$$y(x) = \frac{1}{p} \frac{\beta_y}{1 - \beta_z} (x - h(x)). \quad (25)$$

Market clearing imposes

$$\int_{x_{\min}}^{\infty} y(x) dG_x(x) = \mu_m \int_{y_c}^{\infty} y dG_y(y), \quad (26)$$

where $y(x)$ denotes the consumption of beer by a generalist of income x and G_a the cdf

of variable a . Plugging (25) in (26) we obtain:

$$py_c = \frac{\psi}{\mu_m} \left(\frac{y_c}{y_{\min}} \right)^{\alpha_y},$$

$$\text{with } \psi \equiv \frac{\alpha_y - 1}{\alpha_y} \frac{\alpha_z \beta_y \left(\frac{1}{\alpha_x} + \lambda - 1 \right)}{\lambda (\beta_z + \alpha_z (1 - \beta_z))} \hat{x}.$$

This implies that there are two possible scenarios. If $\psi \geq \mu_m x_{\min}$ then $py_{\min} \geq x_{\min}$ so that all possible brewers end up working as brewers. We then have

$$p = \frac{\psi}{\mu_m y_{\min}}.$$

Since ψ is decreasing in α_x , a decrease in the shape parameter of generalist income is associated with a proportional increase in brewer's income.

Note that

$$\frac{\psi}{x_{\min}} = \frac{\alpha_z \beta_y (\alpha_y - 1)}{\lambda (\beta_z + \alpha_z (1 - \beta_z)) \alpha_y} \frac{1 + (\lambda - 1) \alpha_x}{\alpha_x - 1}$$

is decreasing in α_x . Therefore as α_x decreases then this situation becomes more and more likely.

Otherwise, $y_c > y_{\min}$ with

$$y_c = y_{\min} \left(\mu_m \frac{x_{\min}}{\psi} \right)^{\frac{1}{\alpha_y}},$$

so that as α_x decreases (and consequently x_{\min} to keep mean income of generalists constant), y_c decreases and more and more potential brewers decide to become brewers. This leads to

$$p = \left(\frac{\psi}{\mu_m} \right)^{\frac{1}{\alpha_y}} \left(\frac{\alpha_x - 1}{\alpha_x} \hat{x} \right)^{\frac{\alpha_y - 1}{\alpha_y}} \frac{1}{y_{\min}}$$

$$= \left(\frac{1}{\mu_m} \frac{\alpha_z \beta_y (\alpha_y - 1)}{\lambda (\beta_z + \alpha_z (1 - \beta_z)) \alpha_y} \right)^{\frac{1}{\alpha_y}} \left(\frac{1 + (\lambda - 1) \alpha_x}{\alpha_x - 1} \right)^{\frac{1}{\alpha_y}} \frac{\alpha_x - 1}{\alpha_x} \frac{\hat{x}}{y_{\min}}.$$

Note that

$$\begin{aligned} & \frac{d}{d\alpha_x} (1 + (\lambda - 1) \alpha_x)^{\frac{1}{\alpha_y}} \frac{(\alpha_x - 1)^{\frac{\alpha_y - 1}{\alpha_y}}}{\alpha_x} \\ &= [(1 + (\lambda - 1) \alpha_x) (\alpha_y - 1) - (\alpha_x - 1)] \frac{(1 + (\lambda - 1) \alpha_x)^{\frac{1}{\alpha_y} - 1} (\alpha_x - 1)^{\frac{-1}{\alpha_y}}}{\alpha_y \alpha_x^2}, \end{aligned}$$

the sign of which is ambiguous since λ can be close to 1 and we may have $\alpha_x > \alpha_y$. Therefore in this case, a decrease in α_x increases the supply of beers but as a result the impact on brewers' income is ambiguous.

For any price level \tilde{p} , we can define the real welfare measure similarly as the income which gives the same utility in the market and when the agent is forced to consume (for free) z_c while having y prices at \tilde{p} . That is we now have:

$$u \left(z_c, \frac{1 - \beta_z - \beta_y}{1 - \beta_z} eq(x), \frac{\beta_y}{1 - \beta_z} \frac{eq(x)}{\tilde{p}} \right) = u(z(x), c(x), y(x)).$$

We then obtain:

$$\begin{aligned} eq(x) &= \left(\frac{\tilde{p}}{p} \right)^{\frac{\beta_y}{1 - \beta_z}} \left(\frac{\alpha_z (1 - \beta_z)}{\alpha_z (1 - \beta_z) + \beta_z \alpha_x} x \right. \\ &\quad \left. - \frac{1}{\lambda} \frac{\alpha_z (1 - \beta_z) + \beta_z \alpha_x (1 - \lambda)}{\alpha_z (1 - \beta_z) + \beta_z \alpha_x} x_{\min} \left(\frac{x}{x_{\min}} \right)^{-\frac{\alpha_x \beta_z}{\alpha_z (1 - \beta_z)}} \right) \left(\frac{x}{x_{\min}} \right)^{\frac{\alpha_x \beta_z}{\alpha_z (1 - \beta_z)}}. \end{aligned}$$

Therefore the analysis of real income inequality is the same whether $\beta_y = 0$ or not.

A.5 Deriving (15)

Using that both doctors talents and income are approximately Pareto distributed, we can rewrite (3) as:

$$(m(\bar{z}) / (m(z)))^{\alpha_x} = \frac{\bar{G}_{\bar{z}, z_c}(\bar{z})}{G_x(m(\bar{z}))} \left((\bar{z}/z)^{\alpha_x} + o \left(\left(\frac{\bar{z}}{z} \right)^{\alpha_x} \right) \right) - o \left(\left(\frac{m(\bar{z})}{m(z)} \right)^{\alpha_x} \right).$$

From this we get that $m(z)$ is of the order of $z^{\frac{\alpha_z}{\alpha_x}}$ and therefore

$$m(z) = B z^{\frac{\alpha_z}{\alpha_x}} + o \left(z^{\frac{\alpha_z}{\alpha_x}} \right)$$

with B defined as in the text. We can then rewrite (3) as

$$w'(z)z = \frac{\beta_z}{1-\beta_z} \left(\lambda B z^{\frac{\alpha_z}{\alpha_x}} - w(z) \right) + o\left(z^{\frac{\alpha_z}{\alpha_x}}\right). \quad (27)$$

We then define $\bar{w}(z) \equiv \frac{\beta_z \alpha_x}{\alpha_z(1-\beta_z) + \beta_z \alpha_x} \lambda B z^{\frac{\alpha_z}{\alpha_x}}$ which is a solution to the differential equation without the negligible term, and $\tilde{w}(z) \equiv w(z) - \bar{w}(z)$, which must satisfy

$$\tilde{w}'(z)z = -\frac{\beta_z}{1-\beta_z} \tilde{w}(z) + o\left(z^{\frac{\alpha_z}{\alpha_x}}\right).$$

This gives

$$\tilde{w}'(z)z^{\frac{\beta_z}{1-\beta_z}} + \frac{\beta_z}{1-\beta_z} \tilde{w}(z)z^{\frac{\beta_z}{1-\beta_z}-1} = o\left(z^{\frac{\alpha_z}{\alpha_x}}z^{\frac{\beta_z}{1-\beta_z}-1}\right)$$

Integrating we obtain:

$$\tilde{w}(z) = K z^{-\frac{\beta_z}{1-\beta_z}} + o\left(z^{\frac{\alpha_z}{\alpha_x}}\right)$$

for some constant K , therefore $\tilde{w}(z)$ is negligible in front of $\bar{w}(z)$.

A.6 Proof of Proposition 3

We rewrite (19) more precisely as:

$$w'(z) = \lambda \frac{\varepsilon-1}{\varepsilon} \frac{\beta_z}{\beta_c} z^{-\frac{1}{\varepsilon}} \left(\lambda B z^{\frac{\alpha_z}{\alpha_x}} - w(z) \right)^{\frac{1}{\varepsilon}} + o\left(\frac{\lambda B z^{\frac{\alpha_z}{\alpha_x}} - w(z)}{z}\right)^{\frac{1}{\varepsilon}}. \quad (28)$$

Since consumption of the homogeneous good must remain positive then $\lim \lambda B z^{\frac{\alpha_z}{\alpha_x}} - w(z) \geq 0$, which means that $w(z)$ cannot grow faster than $z^{\frac{\alpha_z}{\alpha_x}}$. We can then distinguish 2 cases: $w(z) = o\left(z^{\frac{\alpha_z}{\alpha_x}}\right)$ and $w(z) \propto z^{\frac{\alpha_z}{\alpha_x}}$.

Case with $w(z) = o\left(z^{\frac{\alpha_z}{\alpha_x}}\right)$. Then for z high enough, one obtains that

$$w'(z) = \lambda \frac{\beta_z}{\beta_c} B^{\frac{1}{\varepsilon}} z^{\left(\frac{\alpha_z}{\alpha_x}-1\right)\frac{1}{\varepsilon}} + o\left(z^{\left(\frac{\alpha_z}{\alpha_x}-1\right)\frac{1}{\varepsilon}}\right). \quad (29)$$

Integrating, we obtain that for $\left(\frac{\alpha_z}{\alpha_x} - 1\right)\frac{1}{\varepsilon} \neq -1$

$$w(z) = K + \lambda \frac{\beta_z}{\beta_c} \frac{B^{\frac{1}{\varepsilon}}}{\left(\frac{\alpha_z}{\alpha_x} - 1\right)\frac{1}{\varepsilon} + 1} z^{\left(\frac{\alpha_z}{\alpha_x}-1\right)\frac{1}{\varepsilon}+1} + o\left(z^{\left(\frac{\alpha_z}{\alpha_x}-1\right)\frac{1}{\varepsilon}+1}\right),$$

where K is a constant. Note that to be consistent, we must have $\left(\frac{\alpha_z}{\alpha_x} - 1\right) \frac{1}{\varepsilon} + 1 < \frac{\alpha_z}{\alpha_x}$, that is $(\alpha_z - \alpha_x)(\varepsilon - 1) > 0$: this case is ruled out if $\alpha_z \geq \alpha_x$ and $\varepsilon < 1$ or if $\alpha_z \leq \alpha_x$ and $\varepsilon > 1$.

If $\left(\frac{\alpha_z}{\alpha_x} - 1\right) \frac{1}{\varepsilon} + 1 < 0$ then $w(z)$ is bounded by K .

If $\left(\frac{\alpha_z}{\alpha_x} - 1\right) \frac{1}{\varepsilon} + 1 > 0$, then we get that

$$w(z) = f^w(z) \equiv \lambda \frac{\beta_z}{\beta_c} \frac{B^{\frac{1}{\varepsilon}}}{\left(\frac{\alpha_z}{\alpha_x} - 1\right) \frac{1}{\varepsilon} + 1} z^{\left(\frac{\alpha_z}{\alpha_x} - 1\right) \frac{1}{\varepsilon} + 1} + o\left(z^{\left(\frac{\alpha_z}{\alpha_x} - 1\right) \frac{1}{\varepsilon} + 1}\right),$$

where the notation f^w is introduced to help notation. Therefore one gets, for \bar{w} large:

$$\Pr(W > w) = \Pr(Z > (f^w)^{-1}(w)) = \bar{G}_w(\bar{w}) \left(\frac{\bar{w}}{w}\right)^{\frac{\alpha_z}{\left(\frac{\alpha_z}{\alpha_x} - 1\right) \frac{1}{\varepsilon} + 1}} + o\left(w^{-\frac{\alpha_z}{\left(\frac{\alpha_z}{\alpha_x} - 1\right) \frac{1}{\varepsilon} + 1}}\right),$$

so that w is Pareto distributed asymptotically with a coefficient $\alpha_w = \frac{\alpha_z}{\left(\frac{\alpha_z}{\alpha_x} - 1\right) \frac{1}{\varepsilon} + 1}$, which is increasing in α_x (and we have $\alpha_w > \alpha_x$).

If $\left(\frac{\alpha_z}{\alpha_x} - 1\right) \frac{1}{\varepsilon} + 1 = 0$, then $\alpha_z = \alpha_x(1 - \varepsilon)$, and integrating (29), one obtains

$$w(z) = f^w(z) \equiv \lambda \frac{\beta_z}{\beta_c} B^{\frac{1}{\varepsilon}} \ln z + o(\ln z).$$

Therefore

$$\begin{aligned} \Pr(W > w) &= \Pr\left(Z > \left(\exp\left(\frac{\beta_c}{\lambda \beta_z B^{\frac{1}{\varepsilon}}} w\right) + o(\exp(w))\right)\right) \\ &= \bar{G}_{z, z_c}(\bar{z}) \bar{z}^{\alpha_z} \exp\left(-\frac{\alpha_z \beta_c}{\lambda \beta_z B^{\frac{1}{\varepsilon}}} w\right) + o(\exp(-\alpha_z w)) \end{aligned}$$

In that case, w is distributed exponentially.

Case where $w(z) \propto z^{\frac{\alpha_z}{\alpha_x}}$. That is we assume that

$$w(z) = Az^{\frac{\alpha_z}{\alpha_x}} + o\left(z^{\frac{\alpha_z}{\alpha_x}}\right) \tag{30}$$

for some constant $A > 0$. Then, we have that

$$\begin{aligned}\Pr(W > w) &= \Pr\left(Z > \left(\left(\frac{w}{A}\right)^{\frac{\alpha_x}{\alpha_z}} + o(w)^{\frac{\alpha_x}{\alpha_z}}\right)\right) \\ &= \bar{G}_w(\bar{w}) \left(\frac{\bar{w}}{w}\right)^{\alpha_x} + o(w)^{\frac{\alpha_x}{\alpha_z}}\end{aligned}$$

That is w is Pareto distributed with coefficient α_x .

Plugging (30) in (28), we get:

$$A \frac{\alpha_z}{\alpha_x} z^{\frac{\alpha_z}{\alpha_x} - 1} + o\left(z^{\frac{\alpha_z}{\alpha_x} - 1}\right) = \lambda \frac{\varepsilon - 1}{\varepsilon} \frac{\beta_z}{\beta_c} (\lambda B - A)^{\frac{1}{\varepsilon}} z^{\left(\frac{\alpha_z}{\alpha_x} - 1\right) \frac{1}{\varepsilon}} + o\left((\lambda B - A)^{\frac{1}{\varepsilon}} z^{\left(\frac{\alpha_z}{\alpha_x} - 1\right) \frac{1}{\varepsilon}}\right). \quad (31)$$

First, assume that $\alpha_z = \alpha_x$, then we get that the solution is characterized by $A = \lambda \frac{\varepsilon - 1}{\varepsilon} \frac{\beta_z}{\beta_c} (\lambda B - A)^{\frac{1}{\varepsilon}}$.

Consider now that $\alpha_z \neq \alpha_x$. If $\lambda B \neq A$ then (31) is impossible when $\varepsilon \neq 1$, therefore we must have that $\lambda B = A$. This equation then requires that

$$\frac{\alpha_z}{\alpha_x} - 1 < \left(\frac{\alpha_z}{\alpha_x} - 1\right) \frac{1}{\varepsilon} \Leftrightarrow (\alpha_z - \alpha_x)(\varepsilon - 1) < 0.$$

In fact, for $(\alpha_z - \alpha_x)(\varepsilon - 1) < 0$, one gets that

$$w(z) = \lambda B z^{\frac{\alpha_z}{\alpha_x}} - \lambda \left(B \frac{\alpha_z}{\alpha_x} \frac{\beta_c}{\beta_z}\right)^{\varepsilon} z^{\varepsilon \left(\frac{\alpha_z}{\alpha_x} - 1\right) + 1} + o\left(z^{\varepsilon \left(\frac{\alpha_z}{\alpha_x} - 1\right) + 1}\right)$$

satisfies (28) provided that the function $o\left(z^{\varepsilon \left(\frac{\alpha_z}{\alpha_x} - 1\right) + 1}\right)$ solves the appropriate differential equation.

Collecting the different cases together gives proposition 3.

A.7 Different adjustment margin

In this appendix we fully solve the model described in section 2.3. First note that above a certain threshold, there will be individuals choosing to be doctors. Assume that this is not the case, then there is an upper bound z_M on the quality of health care provided. Consider an individual 1 with income X who is a generalist. Her utility obeys $u(X) \leq z_M^{\beta_z} X^{1 - \beta_z}$. Consider now individual 2 with generalist ability $X^{\frac{1}{2}}$. For X large enough, this individual would be a generalist. Assume, however, that she switches and decides to become a doctor, then she would provide health care service

with quality $z_{\min} \left(\frac{X^{\frac{1}{2}}}{x_{\min}} \right)^{\frac{\alpha_x}{\alpha_z}}$. Individual 1 would then rather hire individual 2 as a doctor and consume $\frac{1}{2}X$ in homogeneous good. Under this alternative allocation her utility is $\left(\frac{1}{2}\right)^{1-\beta_z} X^{1-\beta_z} z_{\min}^{\beta_z} \left(\frac{X^{\frac{1}{2}}}{x_{\min}} \right)^{\frac{\alpha_x}{\alpha_z} \beta_z}$, which for X high enough is higher than the utility under the original allocation. Individual 2 earns $\frac{1}{2}X$ which is also higher than her initial income. Therefore this is a profitable deviation and the initial allocation cannot be an equilibrium.

As a result, the equilibrium must be that below a certain rank some individuals choose to be doctors. We then have 3 possible cases, which we will solve in turn:

- Below a certain rank individuals choose to be both doctors and generalists and above it they all choose to be generalists;
- Below a certain rank individuals choose to be both doctors and generalists and above it they all choose to be doctors;
- Below a certain rank, all individuals choose to be doctors.

Case 1. Consider first the case where there exists a z_c such that individuals of rank higher than $\bar{G}_z(z_c)$ all choose to be generalists. Then (12) applies for $z > z_c$ and we know that for $z \geq z_c$, $\mu = \frac{\lambda^{\alpha_x - 1}}{\left(\frac{\alpha_z}{\alpha_x} \frac{1-\beta_z}{\beta_z} + 1\right)^{\alpha_x}}$, which we assume to be smaller than 1. Since $m(z_c) = x_{\min}$, we obtain:

$$z_c = z_{\min} \left(\frac{\lambda}{\frac{\alpha_z}{\alpha_x} \frac{1-\beta_z}{\beta_z} + 1} \right)^{\frac{\alpha_x}{\alpha_z}}, \quad (32)$$

which is only possible if $\lambda \geq \frac{\alpha_z}{\alpha_x} \frac{1-\beta_z}{\beta_z} + 1$.

Case 2. Consider now the opposite case. Individuals ranked above $\bar{G}_z(z_m)$ all choose to be generalists, those ranked below are indifferent. Since $\lambda > 1$, the supply of health services by agents ranked higher than $\bar{G}_z(z_m)$ is enough to cover their own demand for health services. Therefore, if one denotes by $r(z)$ the rank of the patient of a doctor of quality z , we obtain that there exists a $z_p < z_m$, such that $r(z_p) = z_m$: doctors with ability lower than z_p only provide health services to doctors and those with ability above z_p provide health services to both doctors and generalists. Since $z_m > z_p$, we have that for $z \geq z_m$, (12) applies which directly leads to $\mu = \frac{\lambda^{\alpha_x - 1}}{\left(\frac{\alpha_z}{\alpha_x} \frac{1-\beta_z}{\beta_z} + 1\right)^{\alpha_x}}$ for $z \geq z_m$. This imposes, as before, the restriction $\frac{\lambda^{\alpha_x - 1}}{\left(\frac{\alpha_z}{\alpha_x} \frac{1-\beta_z}{\beta_z} + 1\right)^{\alpha_x}} < 1$. We then get to further write for

$z \leq z_m$:

$$r(z) = \int_z^{z_m} \lambda g_z(\zeta) d\zeta + \int_{z_m}^{\infty} \lambda \mu(\zeta) g_z(\zeta) d\zeta = \lambda \left(\left(\frac{z_{\min}}{z} \right)^{\alpha_z} - (1-\mu) \left(\frac{z_{\min}}{z_m} \right)^{\alpha_z} \right). \quad (33)$$

For $z \geq z_p$, $m(z) = \overline{G}_x^{-1}(r(z))$, so that (33) implies

$$m(z) = x_{\min} \lambda^{-\frac{1}{\alpha_x}} \left(\left(\frac{z_{\min}}{z} \right)^{\alpha_z} - (1-\mu) \left(\frac{z_{\min}}{z_m} \right)^{\alpha_z} \right)^{-\frac{1}{\alpha_x}} \text{ for } z \in (z_p, z_m).$$

(3) still applies and now gives the differential equation:

$$\left(w'(z) z + \frac{\beta_z}{1-\beta_z} w(z) \right) = \frac{\beta_z}{1-\beta_z} x_{\min} \lambda^{\frac{\alpha_x-1}{\alpha_x}} \left(\left(\frac{z_{\min}}{z} \right)^{\alpha_z} - (1-\mu) \left(\frac{z_{\min}}{z_m} \right)^{\alpha_z} \right)^{-\frac{1}{\alpha_x}}.$$

Using that $w(z_m) = x_{\min} \left(\frac{z_m}{z_{\min}} \right)^{\frac{\alpha_z}{\alpha_x}}$, the solution to this differential equation is then given by:

$$w(z) = z^{-\frac{\beta_z}{1-\beta_z}} x_{\min} (z_{\min})^{-\frac{\alpha_z}{\alpha_x}} \left(\frac{\alpha_z}{z_m^{\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1-\beta_z}}} - \frac{\beta_z}{1-\beta_z} \lambda^{\frac{\alpha_x-1}{\alpha_x}} \int_z^{z_m} \zeta^{-\frac{1-2\beta_z}{1-\beta_z}} (\zeta^{-\alpha_z} - (1-\mu) z_m^{-\alpha_z})^{-\frac{1}{\alpha_x}} d\zeta \right).$$

For this to be an equilibrium, we need to check that $w(z) \geq x_{\min} \left(\frac{z}{z_{\min}} \right)^{\frac{\alpha_z}{\alpha_x}}$, which is the income that a doctor of rank $\overline{G}_z(z)$ would obtain as a generalist. We can rewrite:

$$w(z) - x_{\min} \left(\frac{z}{z_{\min}} \right)^{\frac{\alpha_z}{\alpha_x}} = x_{\min} (z_{\min})^{-\frac{\alpha_z}{\alpha_x}} z^{-\frac{\beta_z}{1-\beta_z}} T(z)$$

with

$$T(z) \equiv z_m^{\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1-\beta_z}} - z^{\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1-\beta_z}} - \frac{\beta_z}{1-\beta_z} \lambda^{\frac{\alpha_x-1}{\alpha_x}} \int_z^{z_m} \zeta^{-\frac{1-2\beta_z}{1-\beta_z}} (\zeta^{-\alpha_z} - (1-\mu) z_m^{-\alpha_z})^{-\frac{1}{\alpha_x}} d\zeta.$$

We get

$$T'(z) = \left(1 - \left(\frac{z^{-\alpha_z} - (1-\mu) z_m^{-\alpha_z}}{\mu z^{-\alpha_z}} \right)^{\frac{1}{\alpha_x}} \right) \frac{\beta_z \lambda^{\frac{\alpha_x-1}{\alpha_x}}}{1-\beta_z} z^{-\frac{1-2\beta_z}{1-\beta_z}} (z^{-\alpha_z} - (1-\mu) z_m^{-\alpha_z})^{-\frac{1}{\alpha_x}}.$$

where we used that

$$\frac{\alpha_z}{\alpha_x} \frac{1-\beta_z}{\beta_z} + 1 = \lambda (\mu \lambda)^{-\frac{1}{\alpha_x}}. \quad (34)$$

Further for $z < z_m$, we get that $z^{-\alpha_z} - (1 - \mu) z_m^{-\alpha_z} > \mu z^{-\alpha_z}$, so that $T'(z) < 0$. Since $T(z_m) = 0$, then we get that $T(z) > 0$ for $z < z_m$, which ensures that $w(z) > x_{\min} \left(\frac{z}{z_{\min}} \right)^{\frac{\alpha_z}{\alpha_x}}$ for $z_p \leq z < z_m$.

Finally, we consider what happens for $z < z_p$. Denote by $d(z)$ the doctor's ability of the individual of rank $r(z)$, then using (33) we get:

$$d(z) = \lambda^{-\frac{1}{\alpha_z}} \left(z^{-\alpha_z} - (1 - \mu) z_m^{-\alpha_z} \right)^{-\frac{1}{\alpha_z}}. \quad (35)$$

To close the market, it must be that $d(z_{\min}) = z_{\min}$, which implies that

$$z_m = z_{\min} \left(\frac{1 - \mu}{1 - \frac{1}{\lambda}} \right)^{\frac{1}{\alpha_z}}. \quad (36)$$

Therefore $z_m > z_{\min}$ is only possible if $\mu < 1/\lambda$, which corresponds to $\lambda < \frac{\alpha_z}{\alpha_x} \frac{1 - \beta_z}{\beta_z} + 1$ (the opposite from case 1).

Further, by definition again, we must have $d(z_p) = z_m$, so that:

$$z_p = \frac{z_m}{\left(1 + \frac{1}{\lambda} - \mu\right)^{\frac{1}{\alpha_z}}} = z_{\min} \left(\frac{1 - \mu}{\left(1 - \frac{1}{\lambda}\right) \left(1 + \frac{1}{\lambda} - \mu\right)} \right)^{\frac{1}{\alpha_z}}. \quad (37)$$

It is direct to verify that for $\mu < 1/\lambda$, $z_{\min} < z_p < z_m$.

Now the patient of the doctor of quality z will have an income given by $w(d(z))$. Therefore (3) gives that for $z \leq z_p$, $w(z)$ must satisfy:

$$w'(z) z = \frac{\beta_z}{1 - \beta_z} (\lambda w(d(z)) - w(z)).$$

Multiply this equation by $z^{\frac{\beta_z}{1 - \beta_z} - 1}$ and integrate over (z, z_p) to obtain that the solution must satisfy:

$$w(z) = \left(w(z_p) z_p^{\frac{\beta_z}{1 - \beta_z}} - \int_z^{z_p} \frac{\beta_z}{1 - \beta_z} \zeta^{\frac{2\beta_z - 1}{1 - \beta_z}} \lambda w(d(\zeta)) d\zeta \right) z^{-\frac{\beta_z}{1 - \beta_z}} \text{ for } z \leq z_p.$$

Once again, we need to verify that $w(z) \geq x_{\min} \left(\frac{z}{z_{\min}} \right)^{\frac{\alpha_z}{\alpha_x}}$ for $z < z_p$. Taking the

difference we can write:

$$\begin{aligned}
w(z) - x_{\min} \left(\frac{z}{z_{\min}} \right)^{\frac{\alpha_z}{\alpha_x}} &= \left[\left(w(z_p) - x_{\min} \left(\frac{z_p}{z_{\min}} \right)^{\frac{\alpha_z}{\alpha_x}} \right) z_p^{\frac{\beta_z}{1-\beta_z}} + \frac{x_{\min}}{z_{\min}^{\frac{\alpha_z}{\alpha_x}}} \left(z_p^{\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1-\beta_z}} - z^{\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1-\beta_z}} \right) \right. \\
&\quad \left. - \int_z^{z_p} \frac{\lambda \beta_z \zeta^{\frac{2\beta_z-1}{1-\beta_z}}}{1-\beta_z} w(d(\zeta)) d\zeta \right] z^{-\frac{\beta_z}{1-\beta_z}}.
\end{aligned}$$

We already know that $w(z_p) > x_{\min} z_{\min}^{-\frac{\alpha_z}{\alpha_x}} z_p^{\frac{\alpha_z}{\alpha_x}}$. Moreover for $\zeta \in (z, z_p)$, $d(\zeta) < z_m$, since $w(z)$ is increasing we get

$$w(d(\zeta)) \leq w(z_m) = x_{\min} \left(\frac{z_m}{z_{\min}} \right)^{\frac{\alpha_z}{\alpha_x}}.$$

Therefore, we get:

$$w(z) - x_{\min} \left(\frac{z}{z_{\min}} \right)^{\frac{\alpha_z}{\alpha_x}} > \frac{x_{\min} z^{-\frac{\beta_z}{1-\beta_z}}}{z_{\min}^{\frac{\alpha_z}{\alpha_x}}} T_2(z).$$

with

$$T_2(z) = \left(z_p^{\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1-\beta_z}} - z^{\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1-\beta_z}} \right) - \lambda z_m^{\frac{\alpha_z}{\alpha_x}} \left(z_p^{\frac{\beta_z}{1-\beta_z}} - z^{\frac{\beta_z}{1-\beta_z}} \right).$$

Differentiating, we get:

$$T_2'(z) = \left(\lambda \frac{\beta_z}{1-\beta_z} z_m^{\frac{\alpha_z}{\alpha_x}} - \left(\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1-\beta_z} \right) z^{\frac{\alpha_z}{\alpha_x}} \right) z^{\frac{\beta_z}{1-\beta_z}-1}.$$

Therefore $T_2'(z)$ has the sign of $\lambda \frac{\beta_z}{1-\beta_z} z_m^{\frac{\alpha_z}{\alpha_x}} - \left(\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1-\beta_z} \right) z^{\frac{\alpha_z}{\alpha_x}}$, which is more likely to be negative for a higher z and can change sign at most once on (z_{\min}, z_p) . Using (37) and (34) we get that

$$T_2'(z_p) = \frac{\beta_z \lambda^{\frac{\alpha_x-1}{\alpha_x}} \mu^{-\frac{1}{\alpha_x}}}{1-\beta_z} z^{\frac{\beta_z}{1-\beta_z}-1} z_p^{\frac{\alpha_z}{\alpha_x}} \left(\left(\left(1 + \frac{1}{\lambda} - \mu \right) \lambda \mu \right)^{\frac{1}{\alpha_x}} - 1 \right).$$

Note that $(1 + \frac{1}{\lambda} - \mu) \lambda \mu = 1 - (1 - \mu)(1 - \lambda \mu)$, since $\lambda \mu < 1$ and $\lambda > 1$ (which implies $\mu < 1$), then we get $(1 + \frac{1}{\lambda} - \mu) \lambda \mu < 1$. Therefore $T_2'(z_p) < 0$, so that over (z_{\min}, z_p)

either T_2 is everywhere decreasing or T_2 is initially increasing and afterwards decreasing. In the former case since $T_2(z_p) > 0$, we directly get that $T_2(z) > 0$ for $z \in (z_{\min}, z_p)$. In the latter case, a necessary and sufficient condition to get $T_2(z) > 0$ over the interval (z_{\min}, z_p) is that $T_2(z_{\min}) > 0$.

Using (36) and (37), we now compute

$$T_2(z_{\min}) = z_{\min}^{\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1-\beta_z}} \cdot \left[\lambda \left(\frac{1-\mu}{1-\frac{1}{\lambda}} \right)^{\frac{1}{\alpha_x}} - 1 - \left(\lambda - \left(\frac{1}{1+\frac{1}{\lambda}-\mu} \right)^{\frac{1}{\alpha_x}} \right) \left(\frac{1-\mu}{1-\frac{1}{\lambda}} \right)^{\frac{1}{\alpha_x}} \left(\frac{1-\mu}{(1-\frac{1}{\lambda})(1+\frac{1}{\lambda}-\mu)} \right)^{\frac{1}{\alpha_z} \frac{\beta_z}{1-\beta_z}} \right]$$

Note that $\lambda - \left(\frac{1}{1+\frac{1}{\lambda}-\mu} \right)^{\frac{1}{\alpha_x}} > 0$ since $\frac{1}{\lambda} > \mu$ and that $\frac{1-\mu}{(1-\frac{1}{\lambda})(1+\frac{1}{\lambda}-\mu)} > 1$ so that $\left(\frac{1-\mu}{(1-\frac{1}{\lambda})(1+\frac{1}{\lambda}-\mu)} \right)^{\frac{1}{\alpha_z} \frac{\beta_z}{1-\beta_z}} > 1$, therefore:

$$\begin{aligned} T_2(z_{\min}) &> z_{\min}^{\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1-\beta_z}} \left[\lambda \left(\frac{1-\mu}{1-\frac{1}{\lambda}} \right)^{\frac{1}{\alpha_x}} - 1 - \left(\lambda - \left(\frac{1}{1+\frac{1}{\lambda}-\mu} \right)^{\frac{1}{\alpha_x}} \right) \left(\frac{1-\mu}{1-\frac{1}{\lambda}} \right)^{\frac{1}{\alpha_x}} \right] \\ &\geq z_{\min}^{\frac{\alpha_z}{\alpha_x} + \frac{\beta_z}{1-\beta_z}} \left[\left(\frac{1-\mu}{(1-\frac{1}{\lambda})(1+\frac{1}{\lambda}-\mu)} \right)^{\frac{1}{\alpha_x}} - 1 \right] \\ &> 0, \end{aligned}$$

since $\frac{1-\mu}{(1-\frac{1}{\lambda})(1+\frac{1}{\lambda}-\mu)} > 1$. This guarantees that we always have $T_2(z) > 0$ over (z_{\min}, z_p) , so that we obtain $w(z) > x_{\min} \left(\frac{z}{z_{\min}} \right)^{\frac{\alpha_z}{\alpha_x}}$ for $z \in (z_{\min}, z_m)$, which ensures that we do have an equilibrium: no doctor of rank higher than $\bar{G}_z(z_m)$ would like to switch and be a generalist.

Case 3. We now consider the case where below a certain rank $\bar{G}_z(z_1)$ all individuals choose to be doctors, while above that rank some individuals choose to be generalists.

Consider a $\delta > 0$ and an individual whose ability as a doctor $z \in (z_1, z_1 + \delta)$. Since $\lambda > 1$, labor market clearing imposes that for δ_1 small enough that individual will cure somebody whose rank is above $\bar{G}_z(z_1)$. Therefore the income of the patient is equal to what he would earn as a generalist (since either he is a generalist or must be indifferent between being a doctor himself or a generalist). We can then write labor market clearing

as:

$$\left(\frac{x_{\min}}{m(z)}\right)^{\alpha_x} = \lambda \left(\frac{z_{\min}}{z}\right)^{\alpha_z},$$

so that

$$m(z) = x_{\min} \lambda^{-\frac{1}{\alpha_x}} \left(\frac{z}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_x}}.$$

Using the first order condition (3), we get

$$w'(z)z + \frac{\beta_z}{1-\beta_z}w(z) = \frac{\beta_z}{1-\beta_z}\lambda^{1-\frac{1}{\alpha_x}}x_{\min}\left(\frac{z}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_x}}.$$

Multiplying on both sides by $z^{\frac{\beta_z}{1-\beta_z}-1}$ and integrate over (z_1, z) to obtain

$$\begin{aligned} \int_{z_1}^z \left(w'(\zeta)\zeta + \frac{\beta_z}{1-\beta_z}w(\zeta) \right) \zeta^{\frac{\beta_z}{1-\beta_z}-1} d\zeta &= \int_{z_1}^z \frac{\beta_z}{1-\beta_z} \lambda^{1-\frac{1}{\alpha_x}} x_{\min} \left(\frac{\zeta}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_x}} \zeta^{\frac{\beta_z}{1-\beta_z}-1} d\zeta \\ \Rightarrow w(z)z^{\frac{\beta_z}{1-\beta_z}} - w(z_1)z_1^{\frac{\beta_z}{1-\beta_z}} &= Mx_{\min} \left(\left(\frac{z}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_x}} z^{\frac{\beta_z}{1-\beta_z}} - \left(\frac{z_1}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_x}} z_1^{\frac{\beta_z}{1-\beta_z}} \right). \end{aligned} \quad (38)$$

where we define $M \equiv \frac{\lambda^{1-\frac{1}{\alpha_x}}}{\frac{\alpha_z}{\alpha_x} \frac{1-\beta_z}{\beta_z} + 1}$. Assume that $M < 1$ then we would get $w(z)z^{\frac{\beta_z}{1-\beta_z}} < x_{\min} \left(\frac{z}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_x}}$, for $z > z_1$, contradicting the fact that this individual chooses to be a doctor. Therefore we must have $M \geq 1$.

Assume now that there is a $\delta > 0$, such that individuals ranked between $\bar{G}_z(z_1)$ and $\bar{G}_z(z_1 - \delta)$ are indifferent between being doctors and being generalist. Consider $z \in (z_1 - \delta, z_1)$, then individual with such ability is indifferent between being a doctor and a generalist so that $w(z) = x_{\min} \left(\frac{z}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_x}}$. Health care market clearing can be written as:

$$\left(\frac{x_{\min}}{m(z)}\right)^{\alpha_x} = \lambda \left(\left(\frac{z_{\min}}{z_1}\right)^{\alpha_z} + \int_z^{z_1} \alpha_z \mu(\zeta) \zeta^{-\alpha_z-1} z_{\min}^{\alpha_z} d\zeta \right).$$

This implies that

$$m(z) = x_{\min} \lambda^{-\frac{1}{\alpha_x}} \left(\left(\frac{z_{\min}}{z_d}\right)^{\alpha_z} + \int_z^{z_1} \alpha_z \mu(\zeta) \zeta^{-\alpha_z-1} z_{\min}^{\alpha_z} d\zeta \right)^{-\frac{1}{\alpha_x}}.$$

Plugging in this expression in the first order condition (3) which still holds we obtain:

$$\left(\frac{1}{z_1}\right)^{\alpha_z} + \int_z^{z_1} \alpha_z \mu(\zeta) \zeta^{-\alpha_z-1} d\zeta = \frac{\lambda^{\alpha_x-1}}{\left(\frac{\alpha_z}{\alpha_x} \frac{1-\beta_z}{\beta_z} + 1\right)^{\alpha_x}} z^{-\alpha_z}.$$

Differentiating with respect to z , one gets that $\mu(z) = M^{\alpha_x}$, since $M \geq 1$, then $\mu(z) \geq 1$. We assumed that $\mu(z) < 1$, therefore there is a contradiction: individuals cannot be indifferent between being generalists and doctors. Instead all individuals choose to be generalists.

Therefore the equilibrium must be such that all individuals rank above $\bar{G}_z(z_1)$ are generalists and all ranked below are doctors. Using market clearing for the whole population, we get

$$1 = \lambda (z_1/z_{\min})^{\alpha_z} \Rightarrow z_1 = \lambda^{\frac{1}{\alpha_z}} z_{\min}.$$

More generally market clearing above z_1 implies that for $z > z_1$, $r(z) = \lambda \left(\frac{z_{\min}}{z}\right)^{\alpha_z}$, where, as before, $r(z)$ denotes the rank of the patient of a doctor of quality z . Define z_2 such that $r(z_2) = z_1$, so that $z_2 = \lambda^{\frac{1}{\alpha_z}} z_1$, that is z_2 is the ability of the doctor who cures patients of doctor's ability z_1 . Then all doctors with ability in (z_1, z_2) will cure generalists, while all doctors with ability higher than z_2 will cure doctors (with ability $d(z) = \lambda^{-\frac{1}{\alpha_z}} z$).

Equation (38) applies on (z_1, z_2) , so that one gets

$$w(z) \geq x_{\min} \left(\frac{z}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_x}}.$$

Doctors do not have an incentive to deviate and become generalists.

For $z \geq z_2$, then we get that doctors cure other doctors with ability $d(z) = \lambda^{-\frac{1}{\alpha_z}} z$, (3) then leads to:

$$w'(z) z + \frac{\beta_z}{1-\beta_z} w(z) = \lambda w \left(\lambda^{-\frac{1}{\alpha_z}} z\right).$$

Define $z_i = \lambda^{\frac{1}{\alpha_z} i} z_{\min}$, and assume that over (z_{i-1}, z_i) , we have that $w(z) \geq x_{\min} \left(\frac{z}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_x}}$ (which is true for $i \in \{1, 2\}$) Then for $z \in (z_i, z_{i+1})$:

$$\int_{z_i}^z \left(w'(\zeta) \zeta + \frac{\beta_z}{1-\beta_z} w(\zeta) \right) \zeta^{\frac{\beta_z}{1-\beta_z}-1} d\zeta = \lambda \int_{z_i}^z w \left(\lambda^{-\frac{1}{\alpha_z}} \zeta \right) \zeta^{\frac{\beta_z}{1-\beta_z}-1} d\zeta.$$

Using that $w\left(\lambda^{-\frac{1}{\alpha_z}} \zeta\right) \geq x_{\min} \left(\frac{\lambda^{-\frac{1}{\alpha_z}} \zeta}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_x}}$ since $\lambda^{-\frac{1}{\alpha_z}} \zeta \in (z_{i-1}, z_i)$, one gets:

$$w(z) z^{\frac{\beta_z}{1-\beta_z}} - w(z_i) z_i^{\frac{\beta_z}{1-\beta_z}} \geq M x_{\min} \left(\left(\frac{z}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_x}} z^{\frac{\beta_z}{1-\beta_z}} - \left(\frac{z_i}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_x}} z_i^{\frac{\beta_z}{1-\beta_z}} \right),$$

as $M \geq 1$, then

$$\begin{aligned} w(z) z^{\frac{\beta_z}{1-\beta_z}} &\geq x_{\min} \left(\frac{z}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_x}} z^{\frac{\beta_z}{1-\beta_z}} + \left(w(z_i) - \left(\frac{z_i}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_x}}\right) z_i^{\frac{\beta_z}{1-\beta_z}}, \\ &\Rightarrow w(z) = x_{\min} \left(\frac{z}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_x}}. \end{aligned}$$

Therefore by recursivity, we get that for all $z \geq z_1$, doctors do not have an incentive to become generalists, which ensures that this is indeed an equilibrium.

Summary. Consider $M = \frac{\lambda^{1-\frac{1}{\alpha_x}}}{\frac{\alpha_z}{\alpha_x} \frac{1-\beta_z}{\beta_z} + 1}$ We have three cases:

- If $M \geq 1$, then individuals rank below $\bar{G}_z \left(\lambda^{\frac{1}{\alpha_z}} z_1\right)$ are all doctors those above are all generalists.
- If $M < 1$ and $\lambda \geq \frac{\alpha_z}{\alpha_x} \frac{1-\beta_z}{\beta_z} + 1$, a fixed share $\mu = M^{\alpha_x}$ choose to be doctors below a certain rank and all choose to be generalists above that rank.
- If $M < 1$ and $\lambda < \frac{\alpha_z}{\alpha_x} \frac{1-\beta_z}{\beta_z} + 1$. a fixed share $\mu = M^{\alpha_x}$ choose to be doctors below a certain rank and all choose to be doctors above that rank.

A.8 Disentangling supply side and demand side effects

To disentangle supply-side and demand side effects in section 2.3, we now build a model where doctors have an outside option positively correlated with their ability but where patients are a separate group. Formally, there are two types of agents: a mass 1 of generalists, with income x distributed with the Pareto distribution $P(X > x) = (x_{\min}/x)^{\alpha_x}$ and a mass M of potential doctors. Generalists consume the homogeneous good and health care services according to the utility function 1. Potential doctors only consume the homogeneous good, as in section 2.3, they are ranked in descending order of ability and we denote i their rank. Agent i can choose between being a doctor providing health services of quality $z(i)$ and earning $w(z(i))$ or working in the homogeneous good sector

earning $y(i)$. y and z are distributed according to the countercumulative distributions:

$$\bar{G}_y(y(i)) = \bar{G}_z(z(i)) = i \text{ with } \bar{G}_y = \left(\frac{y_{\min}}{y}\right)^{\alpha_y} \text{ and } \bar{G}_z = \left(\frac{z_{\min}}{z}\right)^{\alpha_z}.$$

Further $\lambda M > 1$ and $\lambda > 1$ so that everybody can get health services.

Assume that the equilibrium is such that for individuals of a sufficiently high level of ability, some will choose to be doctors and others to work in the homogeneous good sector. That is for i low enough, agents must be indifferent between becoming a doctor or working in the homogeneous good sector, so that we must have $w(z(i)) = y(i)$. Hence, the wage function must satisfy:

$$w(z) = y_{\min} \left(\frac{z}{z_{\min}}\right)^{\frac{\alpha_z}{\alpha_y}}. \quad (39)$$

Market clearing for health care services above z implies:

$$\left(\frac{x_{\min}}{m(z)}\right)^{\alpha_x} = \lambda M \int_z^{\infty} \mu(\zeta) g_z(\zeta) d\zeta, \quad (40)$$

where $\mu(\zeta)$ denotes the share of potential doctors who decide to work as doctors. Hence:

$$m(z) = x_{\min} \left(\int_z^{\infty} \lambda M \mu(\zeta) g_z(\zeta) d\zeta\right)^{-\frac{1}{\alpha_x}}.$$

Plugging this expression in the first order condition (3) together with (39), we obtain:

$$\int_z^{\infty} \mu(\zeta) g_z(\zeta) d\zeta = \frac{1}{\lambda M} \left(\frac{\frac{\beta_z}{1-\beta_z} \lambda x_{\min}}{\left(\frac{\alpha_z}{\alpha_y} + \frac{\beta_z}{1-\beta_z}\right) y_{\min}}\right)^{\alpha_x} \left(\frac{z}{z_{\min}}\right)^{-\alpha_x \frac{\alpha_z}{\alpha_y}}. \quad (41)$$

Taking the derivative with respect to z , we get:

$$\mu(z) = \frac{\alpha_x}{\alpha_y} \frac{1}{\lambda M} \left(\frac{\frac{\beta_z}{1-\beta_z} \lambda x_{\min}}{\left(\frac{\alpha_z}{\alpha_y} + \frac{\beta_z}{1-\beta_z}\right) y_{\min}}\right)^{\alpha_x} \left(\frac{z}{z_{\min}}\right)^{\alpha_z \left(1 - \frac{\alpha_x}{\alpha_y}\right)}. \quad (42)$$

Since $\mu(z) \in (0, 1)$, this case is only possible if $\alpha_y \leq \alpha_x$, that is generalists income distribution has a fatter tail than the outside option for potential doctors (and $\frac{\alpha_x}{\alpha_y} \frac{1}{\lambda M} \left(\frac{\alpha_y \beta_z \lambda x_{\min}}{(\alpha_z(1-\beta_z) + \beta_z \alpha_y) y_{\min}}\right)^{\alpha_x} \leq 1$ if $\alpha_y = \alpha_x$). We then obtain that doctor's income

distribution obeys (for w high enough):

$$\Pr(W > w) = \int_{z_{\min} \left(\frac{w}{y_{\min}}\right)^{\frac{\alpha_y}{\alpha_z}}}^{\infty} \mu(\zeta) \left(\frac{z_{\min}}{\zeta}\right)^{\alpha_z} \frac{d\zeta}{\zeta} = \frac{1}{\lambda M \alpha_z} \left(\frac{\alpha_y \beta_z \lambda x_{\min}}{\alpha_z (1 - \beta_z) + \beta_z \alpha_y}\right)^{\alpha_x} w^{-\alpha_x}.$$

Therefore doctors' income is distributed like the patients' income and not according to doctors' outside option.

With $\alpha_y > \alpha_x$ or $\alpha_y = \alpha_x$ together with $\frac{\alpha_x}{\alpha_y} \frac{1}{\lambda M} \left(\frac{\alpha_y \beta_z \lambda x_{\min}}{(\alpha_z (1 - \beta_z) + \beta_z \alpha_y) y_{\min}}\right)^{\alpha_x} > 1$, then above a certain threshold, all potential doctors will choose to be doctors, so that the model behaves like that of section 2.1.

Therefore, in all cases, at the top, income is distributed in a Pareto way with shape parameter α_x . Changes in α_y have no impact on doctors' top income inequality.

A.9 Proof of Proposition 2

Since $\omega(z)$ is equalized between the two regions, then the threshold z_c of the least able potential doctor must also be the same in the two regions.²⁹ Summing up the market clearing equations (13) and (14), we obtain that as in the baseline model, $z_c = (\lambda \mu_d)^{\frac{1}{\alpha_z}} z_{\min}$. Next combining (13) and (14), we get that

$$x_{\min}^A (1 + \varphi(z))^{-\frac{1}{\alpha_x^A}} = x_{\min}^B \left(\frac{z}{z_c}\right)^{\frac{\alpha_z}{\alpha_x^B} - \frac{\alpha_z}{\alpha_x^A}} (1 - \varphi(z))^{-\frac{1}{\alpha_x^B}}. \quad (43)$$

Since $\alpha_x^B > \alpha_x^A$, we get that $\left(\frac{z}{z_c}\right)^{\frac{\alpha_z}{\alpha_x^B} - \frac{\alpha_z}{\alpha_x^A}}$ tends towards 0. As a net share $\varphi(z) \in (-1, 1)$, if $\varphi(z) \rightarrow -1$, then the left-hand side would tend toward infinity and the right-hand side toward 0, which is a contradiction. Therefore $1 + \varphi(z)$ must be bounded below, which ensures that the left-hand side is bounded above 0. If $\varphi(z) \nearrow 1$, then the right-hand side would be asymptotically 0, this is also a contradiction. Therefore asymptotically, we must have that $\varphi(z) \rightarrow 1$: nearly all the best doctors move to the most unequal region.

Plugging (13) in (3), we get that in region A:

$$w'(z) z + \frac{\beta_z}{1 - \beta_z} w(z) = \frac{\beta_z \lambda}{1 - \beta_z} (1 + \varphi(z))^{-\frac{1}{\alpha_x^A}} \left(\frac{z_c}{z}\right)^{-\frac{\alpha_z}{\alpha_x^A}}.$$

²⁹Here potential doctors who decide to work in the homogeneous good sector would go to region B since $\alpha_x^A > \alpha_x^B$ implies that $x_{\min}^A < x_{\min}^B$. This is without consequences: alternatively, we could have assumed that the outside option of doctors is to produce \hat{x} , which is identical between the two regions. In that case potential doctors who work in the homogeneous sector would not move.

Therefore, asymptotically:

$$w(z) \rightarrow \frac{\lambda \beta_z \alpha_x^A 2^{-\frac{1}{\alpha_x^A}}}{\alpha_z (1 - \beta_z) + \beta_z \alpha_x^A} \left(\frac{z}{z_c} \right)^{\frac{\alpha_z}{\alpha_x^A}} \quad (44)$$

Since $\varphi(z) \rightarrow 1$, after the location decision, doctors' talent is asymptotically distributed with Pareto coefficient α_z in region A : for z high enough, there are $2\mu_d (z_{\min}/z)^{\alpha_z}$ doctors eventually located in region A . We then directly get that doctor's income distribution is asymptotically Pareto distributed with coefficient α_x^A .

From (43), we get that:

$$\begin{aligned} 1 - \varphi(z) &= (x_{\min}^B/x_{\min}^A)^{\alpha_x^B} (1 + \varphi(z))^{\alpha_x^B/\alpha_x^A} (z/z_c)^{\alpha_z(1-\alpha_x^B/\alpha_x^A)} \\ &\rightarrow 2^{\alpha_x^B/\alpha_x^A} (x_{\min}^B/x_{\min}^A)^{\alpha_x^B} (z/z_c)^{\alpha_z(1-\alpha_x^B/\alpha_x^A)}. \end{aligned} \quad (45)$$

Then we can write that in region B , the probability that a doctor earns at least \tilde{w} is given by:

$$P_{doc}^B(W > \tilde{w}) = \frac{\mu_d P(Z > w^{-1}(\tilde{w})) (1 - \varphi(w^{-1}(\tilde{w})))}{\mu_d P(Z > z_c)},$$

where w above denotes the wage function. Indeed, there are originally $\mu_d P(Z > w^{-1}(\tilde{w}))$ doctors present in region B with a talent sufficient to earn \tilde{w} . Out of these doctors, $1 - \varphi(w^{-1}(\tilde{w}))$ stay in region B . Moreover, the total mass of active doctors in region B is given by $\mu_d P(Z > z_c)$, since overall there is no net movement of actual doctors. Using (44) we get that,

$$w^{-1}(\tilde{w}) \rightarrow z_c \left(\tilde{w}^{\alpha_z} \frac{\alpha_z (1 - \beta_z) + \beta_z \alpha_x^A}{\lambda \beta_z \alpha_x^A} 2^{\frac{1}{\alpha_x^A}} \right)^{\frac{\alpha_x^A}{\alpha_z}}.$$

Using this expression and (45) we get that:

$$P_{doc}^B(W > \tilde{w}) = \left(\frac{z_c}{w^{-1}(\tilde{w})} \right)^{\alpha_z} (1 - \varphi(w^{-1}(\tilde{w}))) \rightarrow \left(\frac{x_{\min}^B}{x_{\min}^A} \frac{\lambda \beta_z \alpha_x^A}{\alpha_z (1 - \beta_z) + \beta_z \alpha_x^A} \frac{1}{\tilde{w}} \right)^{\alpha_x^B}.$$

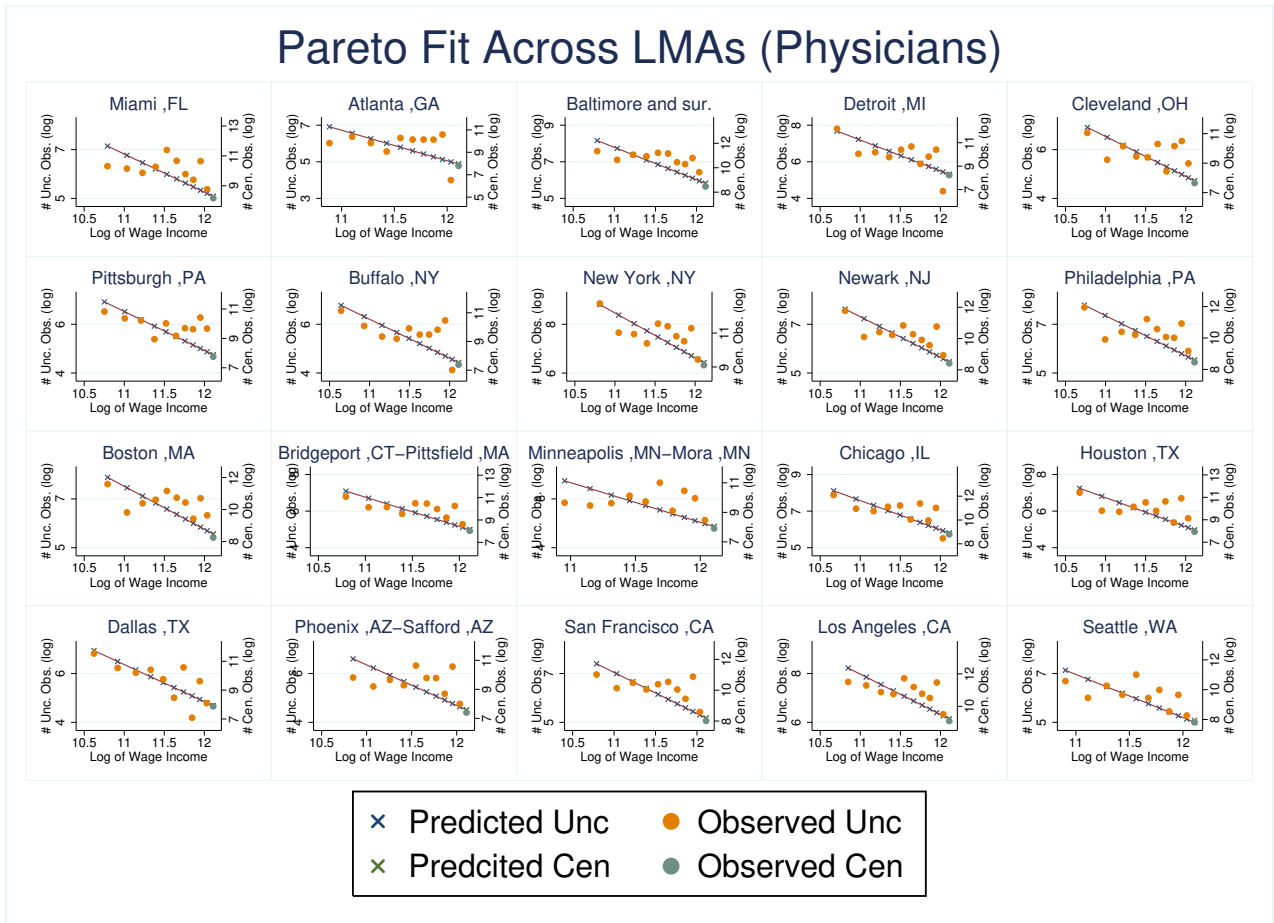
This establishes Proposition 2.

B Pareto Fit and Tables for Top Occupations

Table B.1 gives the change in α for the top occupations. The top occupations for 1980 and 2014 are given in Table B.2.

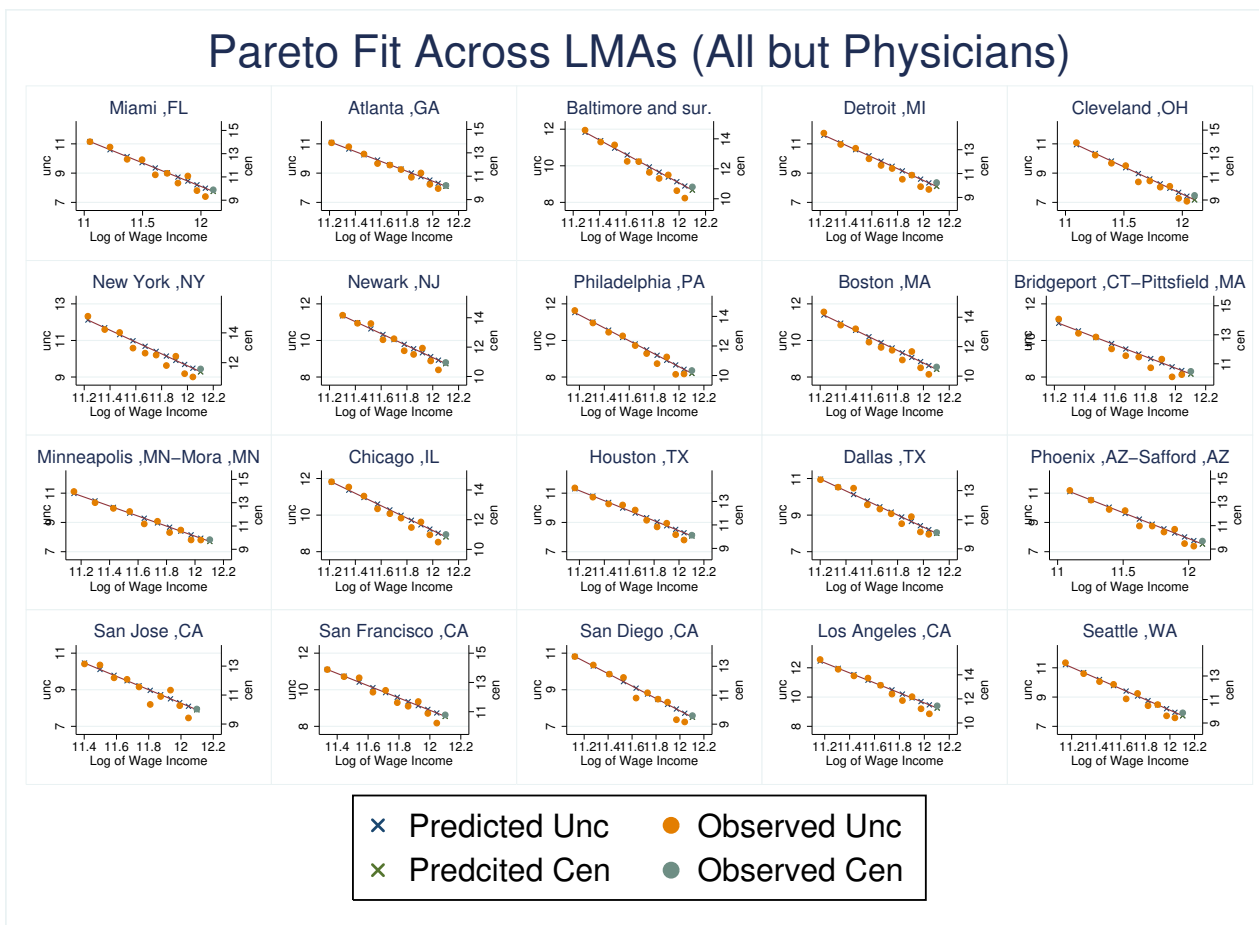
The paper uses the assumption of Pareto for physicians on the LMA-year-occupation level, for LMA-year for the general population and for occupation-year level for the top 20 occupations. Figure 3 in the main text shows the fit with Pareto distribution for the biggest LMA for the whole distribution and for physicians specifically. Figures B.1 and B.2 show analogous figures for the 20 biggest labor market areas for physicians and for the all other occupations than physicians both for the year 2000. Whereas the general population fits the Pareto assumption remarkably well, there is more noise around the line for the physicians, though no systematic deviation.

Figure B.1: Fit to the Pareto Distribution for general income distribution for Physicians for 20 biggest labor market areas for 2000 (using top 65 per cent of uncensored observations)



Notes: Using top 65 per cent of uncensored observations

Figure B.2: Fit to the Pareto Distribution for general income distribution excluding Physicians for 20 biggest labor market areas for 2000 (using top 10 per cent of uncensored observations)



Notes: Using top 10 per cent of uncensored observations

Table B.1: Top occupations and income inequality ($1/\alpha$)

Occupation	1980 (pred. 95/90)	$1/\alpha$		2014 (pred. 95/90)
		1990	2000	
Chief executives and public administrators	0.24 (1.18)	0.34	0.65	0.57 (1.48)
Financial managers	0.32 (1.25)	0.43	0.48	0.52 (1.44)
Managers and specialists in marketing, advertising, and public relations	0.30 (1.23)	0.33	0.36	0.37 (1.29)
Managers in education and related fields	0.19 (1.14)	0.24	0.23	0.29 (1.22)
Managers and administrators, n.e.c.	0.43 (1.34)	0.45	0.36	0.38 (1.30)
Accountants and auditors	0.27 (1.21)	0.32	0.38	0.44 (1.35)
Not-elsewhere-classified engineers	0.22 (1.16)	0.23	0.24	0.23 (1.17)
Computer systems analysts and computer scientists	0.16 (1.12)	0.21	0.25	0.25 (1.19)
Physicians	0.47 (1.39)	0.78	0.55	0.62 (1.54)
Registered nurses	0.17 (1.13)	0.17	0.20	0.23 (1.17)
Subject instructors (HS/college)	0.20 (1.14)	0.24	0.28	0.33 (1.26)
Lawyers	0.42 (1.34)	0.53	0.58	0.58 (1.49)
Computer software developers	0.18 (1.13)	0.19	0.23	0.24 (1.18)
Supervisors and proprietors of sales jobs	0.40 (1.32)	0.45	0.44	0.44 (1.36)
Insurance sales occupations	0.42 (1.34)	0.50	0.52	0.58 (1.49)
Salespersons, n.e.c.	0.35 (1.27)	0.40	0.39	0.42 (1.34)
Supervisors of construction work	0.30 (1.23)	0.33	0.29	0.30 (1.23)
Production supervisors or foremen	0.20 (1.15)	0.20	0.26	0.29 (1.23)
Truck, delivery, and tractor drivers	0.20 (1.15)	0.22	0.24	0.26 (1.20)
Military	0.28 (1.21)	0.25	0.28	0.25 (1.19)

Notes: Estimates of $1/\alpha$ for top 20 occupations using top 10 per cent of population

Table B.2: Top occupations in 1%, 5%, 10%, for year 1980 and 2014

rank	1980			2014		
	top 10 pct	top 5 pct	top 1 pct	top 10 pct	top 5 pct	top 1 pct
1	Managers and administrators, n.e.c.	Managers and administrators, n.e.c.	Managers and administrators, n.e.c.	Managers and administrators, n.e.c.	Managers and administrators, n.e.c.	Physicians
2	Salespersons, n.e.c.	Salespersons, n.e.c.	Physicians	Chief executives and public administrators	Physicians	Managers and administrators, n.e.c.
3	Production supervisors or foremen	Production supervisors or foremen	Salespersons, n.e.c.	Computer software developers	Chief executives and public administrators	Chief executives and public administrators
4	Truck, delivery, and tractor drivers	Managers and specialists in marketing, advertising, and public relations	Managers and specialists in marketing, advertising, and public relations	Physicians	Lawyers	Lawyers
5	Managers and specialists in marketing, advertising, and public relations	Physicians	Lawyers	Lawyers	Salespersons, n.e.c.	Supervisors and proprietors of sales jobs
6	Supervisors of construction work	Supervisors and proprietors of sales jobs	Supervisors and proprietors of sales jobs	Salespersons, n.e.c.	Supervisors and proprietors of sales jobs	Salespersons, n.e.c.
7	Supervisors and proprietors of sales jobs	Truck, delivery, and tractor drivers	Insurance sales occupations	Supervisors and proprietors of sales jobs	Computer software developers	Financial managers
8	Physicians	Lawyers	Production supervisors or foremen	Computer systems analysts and computer scientists	Managers and specialists in marketing, advertising, and public relations	Other financial specialists
9	Accountants and auditors	Supervisors of construction work	Real estate sales occupations	Registered nurses	Financial managers	Managers and specialists in marketing, advertising, and public relations
10	Electrical engineer	Financial managers	Airplane pilots and navigators	Accountants and auditors	Accountants and auditors	Accountants and auditors

Notes:

C Empirical Appendix

C.1 Additional Regressions for other occupations

We perform an analysis like that of Tables 5 and 6 for nurses, College professors and Real Estate agents (occupation code 254). Real Estate agents are censored at around top 7 per cent and we use top 20 per cent uncensored observations.

Finally, we show that income inequality for chief executives and public administrators positively predict the income inequality for secretaries in Table C.3.

C.2 Robustness Checks for Physicians

We perform robustness checks for the the regression in Table 5. In particular, Table C.1 shows the regression for different cut-offs. The parameter estimate is generally not far from 1 and remains significant at the 10% level throughout the regressions. Table C.2 shows that the choice of how many LMAs to include does not affect the parameter estimate much.

Table C.1: IV Regressions for Physicians for different cut-offs of Pareto Distribution

cut-off	(1) 35	(2) 40	(3) 45	(4) 55	(5) 65	(6) 75
log(1/α(-o))	1.30*** [0.25, 2.42]	1.48** [0.06, 2.23]	1.41* [-0.35, 2.86]	1.07 [-0.44, 2.00]	1.17* [-0.12, 2.41]	1.82 [-0.34, 3.67]
logpop	0.09 [-0.09, 0.22]	0.13 [-0.11, 0.26]	0.17* [-0.02, 0.34]	0.17* [-0.02, 0.35]	0.20** [0.02, 0.47]	0.31** [0.02, 0.48]
log(income)	0.06 [-0.11, 0.17]	-0.03 [-0.23, 0.12]	-0.25*** [-0.45,-0.05]	-0.31*** [-0.56,-0.17]	-0.32*** [-0.66,-0.16]	-0.25** [-0.71,-0.06]
1990	-0.07 [-0.19, 0.11]	-0.04 [-0.14, 0.15]	-0.20 [-0.70, 0.03]	-0.38*** [-0.92,-0.21]	-0.41*** [-0.97,-0.23]	0.20** [0.08, 0.60]
2000	-0.20 [-0.40, 0.12]	-0.17 [-0.33, 0.19]	-0.08 [-0.36, 0.06]	-0.17** [-0.45,-0.08]	-0.16** [-0.45,-0.04]	0.00 [-0.17, 0.66]
2014	-0.14 [-0.37, 0.22]	-0.09 [-0.27, 0.31]	-0.13*** [-0.19,-0.10]	-0.21*** [-0.28,-0.18]	-0.25*** [-0.34,-0.21]	0.18* [-0.03, 0.94]
Observations	1,012	1,012	1,012	1,012	1,011	1,011

Bootstrapped standard errors based on 100 draws, stratified at the occupation/year/labor market level. 95 pct confidence interval in square parentheses. Income is average wage income for those with positive income o refers to occupation of interest * $p \leq 0.10$, ** $p \leq 0.05$, *** $p \leq 0.01$

Table C.2: IV Regressions for Physicians for different number of LMAs

	(1)	(2)	(3)	(4)	(5)	(6)
LMAs	100	150	200	253	300	All
$\log(1/\alpha(-o))$	1.70*** [0.76, 2.45]	1.35** [0.27, 2.01]	0.83* [-0.16, 1.57]	1.30*** [0.25, 2.42]	1.21** [0.09, 2.10]	1.11** [0.05, 1.87]
logpop	0.17 [-0.03, 0.33]	0.07 [-0.12, 0.20]	0.02 [-0.14, 0.17]	0.09 [-0.09, 0.22]	0.12 [-0.05, 0.21]	0.09 [-0.05, 0.21]
log(income)	-0.04 [-0.22, 0.15]	0.03 [-0.16, 0.17]	0.09 [-0.08, 0.19]	0.06 [-0.11, 0.17]	0.07 [-0.09, 0.22]	0.08 [-0.11, 0.24]
1990	0.16 [-0.22, 0.39]	0.13 [-0.25, 0.26]	0.05 [-0.30, 0.24]	-0.07 [-0.19, 0.11]	0.14 [-0.26, 0.32]	0.11 [-0.24, 0.26]
2000	0.09 [-0.11, 0.22]	0.07 [-0.13, 0.14]	0.02 [-0.16, 0.11]	-0.20 [-0.40, 0.12]	0.07 [-0.14, 0.18]	0.05 [-0.12, 0.15]
2014	-0.02 [-0.10, 0.03]	-0.04** [-0.10,-0.01]	-0.05*** [-0.11,-0.02]	-0.14 [-0.37, 0.22]	-0.06*** [-0.13,-0.03]	-0.08*** [-0.13,-0.04]
Observations	400	600	800	1,012	1,200	1,573

Bootstrapped standard errors based on 100 draws, stratified at the occupation/year/labor market level. 95 pct confidence interval in square parentheses. Income is average wage income for those with positive income
o refers to occupation of interest * $p \leq 0.10$, ** $p \leq 0.05$, *** $p \leq 0.01$

Table C.3: OLS regressions for secretaries on Chief executives and public administrators for 2000 and 2014

	(1)	(2)	(3)	(4)
	Secretaries	Secretaries	Secretaries	Secretaries
Chief executives and public administrators	0.178*** (4.24)	0.177*** (4.28)	0.136** (2.02)	0.136** (2.00)
2014		-0.0953*** (-4.50)	-0.0968*** (-4.64)	-0.0892 (-1.21)
Log of Inc.				-0.0446 (-0.18)
Log of Pop.				0.0391 (0.19)
Observations	769	769	769	769
Adjusted R^2	0.022	0.046	0.095	0.090

t statistics in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Notes: Regressions limited to 2000 and 2014 due to insufficient information on CEOs in 1980 and 1990. Weighted by number of secretaries by LMA. For 8 observations or more. Column (I) is univariate OLS, Column (II) includes time dummy for 2014, Column (III) further includes labor market area fixed effects and Column (IV) controls for average wage income as well as population.

D Construction of Data on Labor Market Areas

The publicly available data from IPUMS gives information on “country group” in 1980 and “Public Use Microdata Area” (PUMA) for 1990 and onward. We wish to assign these to labor market areas. Dorn (2009) uses a probabilistic approach using the aggregate correspondence between county groups/PUMAs and counties and counties and commuting zones and creates a “crosswalk” assigning weights for each country group in 1980 to 1990 commuting zones and for each PUMA to 1990 commuting zones. If a given county group or PUMA is assigned to multiple commuting zones we “split” all individuals in the county group or PUMA and give each weights from the crosswalk. The IPUMS data from 2012 onward uses the PUMA2010 (updated from the 2010 federal census) and we construct a new crosswalk along the same lines as Dorn (2009). Counties are very stable across town and we manually correct for county changes between 2000 and 2010. Finally, since our unit of analysis is labor market areas we use Missouri Census Data Center (<http://mcdc.missouri.edu/websas/geocorr2k.html>) to aggregate commuting zones into labor market areas. Each commuting zone is uniquely assigned to a labor market area. If a single individual had been split into two commuting zones within the same labor market area using Dorn’s algorithm we combine the two into one observation aggregating their weights. Figure D.1 shows the labor market areas for 1990.

Labor market areas for the United States, 1990



Figure D.1: Labor Market Areas as defined for 1990