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QUALITY OF THE BUSINESS ENVIRONMENT VERSUS QUALITY OF LIFE IN A DYNAMIC MODEL OF URBAN COMPOSITION AND GROWTH

DO FIRMS AND HOUSEHOLDS LIKE THE SAME CITIES?

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Quality of the Business Environment Versus Quality of Life in a Dynamic Model of Urban Composition and Growth

Do Firms and Households Like the Same Cities?

by

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ABSTRACT

Appropriately constructed measures of the quality of life and the quality of the business environment should be important determinants of the growth and composition of population across urban areas. This paper examines that question by extending theoretical measures of household quality of life to construct the first ever measure of the quality of the business environment – the value that firms place on the basket of amenities in a metropolitan area. An annual panel of quality of life and quality of business environment values for 37 cities in the United States is then constructed for the 1977 to 1995 period.

A key finding is that many cities attractive to firms are unattractive to households, and vice versa. In addition, estimates from an error correction model (ECM) indicate that improvements in the quality of the business environment and quality of life have strong positive effects on equilibrium city shares of workers, but negative effects on equilibrium city shares of retirees. The former result reflects outward shifts of labor supply and demand in response to improved amenities. The latter is because retirees avoid high land rents that arise from the in-migration of firms and workers. Moreover, following a shock that creates migratory pressures in the system of cities, worker-population shares converge back to long run equilibrium in 8-1/2 years, while retiree-population shares converge back in 6 years. The longer response time of the worker population likely reflects the cost of adjusting the spatial distribution of industry-specific human and physical capital in a coordinated manner.

1. Introduction

In October 1999, Money Magazine rated San Francisco as the best city in which to live in the United States. A few years earlier, Places Rated Almanac gave that distinction to Pittsburgh, a city once known for its aging steel industry and poor air quality. Analogous rankings have also been published on the best places to do business. In May 2000, Forbes ranked Austin, Texas as the city with the best business environment in the United States, while Syracuse, New York was ranked far behind, at 172. Do these rankings suggest that households and firms favor different cities? If so, what are the implications for the growth and character of individual metropolitan areas? Moreover, given the ad hoc nature of popular rankings of urban areas, to what extent can economic theory be used to develop more reliable and more easily interpretable city rankings?

Cities develop because households and firms want to live and do business in those places. In that regard, one would expect city size and growth to increase with the value that households and firms place on a city's amenities. This suggests that appropriately constructed measures of the quality of life and the quality of the business environment should be important determinants of the distribution and composition of population across urban areas. This paper seeks to examine these issues by drawing on two seemingly disparate literatures that are in fact closely linked, the literature on urban quality of life and the literature on the migration decisions of retirees. As will become apparent, in addressing the impact of quality of life and quality of the business environment on metropolitan area growth and character, we also make several important methodological contributions to these literatures as well.

With the aging of the baby boomers, the growth and demographic composition of different cities is becoming increasingly sensitive to the locational preferences of retirees.[1] That view is implicit in many recent state and local initiatives to attract retiree populations. A number of states, for example, have

[1] In 1990, for example, the population 65 and older in the United States was close to 30 million. That number will increase as the baby boomers enter retirement.

developed marketing programs designed to advertise their amenities to recent retirees [Fagan (1988), Stallman and Siegle (1995), Wilkinson (1995)]. In addition, many states have enacted tax policies designed to attract and retain retirees [Stockbridge-Pratt (1997)].[2] These policy efforts are mirrored in an academic literature that examines retiree preferences for local attributes, including sunshine, recreational opportunities, crime, and medical facilities.[3] Most studies obtain results consistent with widely held priors with regard to such amenities; however, the estimated effects of fiscal variables on retiree location choice are mixed and often inconclusive.[4] In the most recent study, Duncombe, Robbins, and Wolf (2000) conclude that reductions in inheritance taxes, income taxes, and property taxes serve to attract retirees, but that the effect of these tax instruments on retiree migration is small.

In contrast to this literature, Graves and Knapp (1988) argue that because retirees largely do not participate in labor markets, they tend to be drawn to attractive cities in which local amenities are capitalized primarily into lower wages rather than higher house rents. This suggests that retirees should

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[2]Over 20 states, for example, have substantially reduced their estate and gift takes since the mid-1980s in order to reduce out-migration of the elderly. See, for example, Eckl (1986) and Dresher (1993), and related commentary by Mackey (1995a and 1995b).

[3] In public policy and academic circles, the question of where retirees locate is not without controversy. Some have argued that migrating retirees may impose a burden on already stretched local public resources (e.g. Bryant and El-Attar (1984), Longino and Biggar (1981), Duncombe, Robbins, and Wolf (2000)). Others have emphasized that recent retirees that migrate to new areas can enhance the local economy and tax base without imposing undue demands on local government services (Crown (1988), Longino and Crown (1989), Sastry (1992)). Nevertheless, although debate continues, there appears to be widespread belief in government circles that attracting and retaining retirees is desirable, at least if the actions of numerous state and local legislatures are any indication.

[4]See Duncombe, Robbins, and Wolf (2000) for a careful review of much of the literature in this area. In addition, findings regarding proximity to medical facilities are mixed. Proximity to hospitals attracts retirees but concentrations of nursing homes repels retirees. The explanation offered by Duncombe, Robbins, and Wolf (2000) is that most retiree moves are best thought of in three stages as discussed by Litwak and Longino (1987). Moves soon after retirement are sensitive to local amenities relative to proximity to family, and may even entail moving further away from adult children (Clark and Wolf (1992). In the mid- to late seventies, often prompted by declining health or the death of a spouse, retirees that move tend to relocate closer to family (e.g. Meyer (1987), Longino and Serow (1992)). Some retirees ultimately make a third move to a nursing home when full-time care becomes necessary. Such moves typically do not cross county lines. In the context of the discussion above, much of the debate about retiree location decisions as relates to government policy is centered on the first stage moves shortly after retirement.

seek out high quality-of-life and low cost-of-living areas. That idea, while compelling, has gone largely untested. In the case of the retirement literature, this may reflect that the theory offered by Graves and Knapp (1988) requires that one take into account the locational preferences of firms, a seemingly unnecessary complication considering the non-working status of retirees. The quality of life literature, on the other hand, is based on an open city model in which the migration of workers and firms drive house rents, wages, and the spatial distribution of employment across cities [e.g. Roback (1982), Blomquist et al (1988), and Gyourko and Tracy (1991)]. That literature, however, has never measured firm valuations of urban amenities, focusing instead on household rankings of urban quality of life.[5]

Building off Graves and Knapp (1988), this paper emphasizes that both households *and* firms are consumers of city-specific amenities. However, because households and firms differ in their objectives – utility maximization versus profit maximization – they likely differ as well in their valuation of a given basket of metropolitan area attributes (denoted Q_H for households and Q_F for firms). Moreover, changes in Q_F shift the labor demand curve in a city while changes in Q_H shift the labor supply curve. Together, Q_F and Q_H are shown to determine land rents and wages for individual metropolitan areas, while simultaneously driving the equilibrium distribution of workers across cities, and the spatial distribution of retirees. Thus, the degree to which firms want to locate in a city has important implications for the location decisions of both workers and retirees. In the case of workers, this occurs because workers and firms must locate together. In the case of retirees, this occurs because workers, firms, and retirees compete for space.

To extend our approach to the data, two important limitations of the existing literature on quality of life must be addressed. First and most important, we construct and estimate a measure for Q_F that is

[5]A rapidly growing empirical literature on agglomeration economies has also made abundantly clear the degree to which different patterns of employment agglomeration affect productivity and growth (Glaeser et al. (1992), Ellison and Glaeser (1997), Henderson et al. (1995), Henderson and Black (1999), Rosenthal and Strange (2000)). That literature, however, largely does not address the influence of quality of life on the city-specific supply of labor.

solidly grounded in economic theory. To our knowledge, this is the first time that such a measure has been developed. This allows us to rank cities on the basis of firm amenity preferences, and ultimately, to evaluate whether firms and households favor similar or different metropolitan areas.

Second, the urban quality of life literature has been static in nature, precluding analysis of the impact of changes in city valuations on the growth and composition of urban areas.[6] In this paper, we draw on 19 years of data from the American Housing Survey (AHS) and Consumer Population Survey (CPS).[7] Applying panel data methods to these data, we are able to greatly simplify construction of Q_H and Q_F relative to prior studies. This allows us to construct the first ever panel of Q_H and Q_F values for 37 U.S. cities. The panel covers the 1977 to 1995 period on an annual basis and enables us to examine the dynamics of worker and retiree population changes across cities in response to changes over time in the quality of life and the quality of the business environment. A striking result from our analysis is that many of the cities least attractive to households are attractive to firms. Detroit, Gary, Baltimore, Philadelphia, and Washington D.C, for example, are all ranked in the bottom quartile in the eyes of households but are all ranked in the upper half by firms.[8] Those differences reflect the very different goals of the two groups – utility maximization versus profit maximization.

A second important set of results concerns tests of the model. Prior studies have emphasized that individual amenities could have opposite or similar appeal to workers and firms, and therefore, ambiguous effects on the equilibrium distribution of population among cities. That emphasis has largely precluded testing of the model [e.g. Blomquist et al (1988)]. In contrast, by analyzing the time series relationship

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[6]A recent exception of sorts is Kahn (1995). However, Kahn (1995) examined quality of life differentials for only four cities in the United States and for just two time periods, 1980 and 1990.

[7]In contrast, Blomquist et al (1988) and Gyouko and Tracy (1991) both use the 1980 Decenial Census in their quality of life studies. Relative to the Census, the AHS has much superior information on housing attributes while the CPS has much superior information on characteristics of workers.

between population shares across cities, quality of the business environment, and quality of life, we are able to test several fundamental predictions of the model.

Most important, we find that improvements in the quality of life (Q_H) and in the quality of the business environment (Q_F) have strong positive effects on urban growth as measured by equilibrium city shares of workers. This suggests that firms and workers seek out cities with attractive amenities, *ceteris paribus*, a central assumption upon which the quality of life literature is based. In sharp contrast, we also find that a city's share of the total population of retirees *falls* as the city becomes more attractive to workers and firms. In addition, the city's ratio of retirees to workers increases with Q_H but declines with an increase in the cost of housing. These results suggest that retirees shy away from cities that appeal to firms and workers in order to avoid the higher land rents that result from increased competition for space, consistent with the theoretical arguments of Graves and Knapp (1988) noted above. An intriguing policy implication follows: government policies designed to attract workers and firms – such as tax breaks offered to industry – may have the unintended effect of repelling retirees.

A final insight concerns the number of years that must elapse before the distribution of retirees and workers across cities adjusts to policy shifts or to other unexpected events that create migratory pressures. Following an unexpected event that creates pressure for net migration, we find that roughly 90 percent of the adjustment of worker-population shares across cities occurs in 8-1/2 years, while retiree-population shares adjust in 6 years. The longer response time of worker populations likely arises because net migration of workers requires coordinated movement of human and physical industry-specific capital, raising the cost of worker migration. Because such costs are largely not relevant for retirees, the distribution of retirees across cities adjusts to unexpected events more quickly.

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[8]Detroit, for example, was ranked last by households but was ranked in the upper quartile by firms. Washington D.C was ranked 13th from the bottom by households but was ranked fourth from the top by firms.

To clarify these and other results, the following section presents the theory that underlies our measures of Q_H and Q_F. Section 3 describes the data while Section 4 describes the Q_H and Q_F estimates. Section 5 presents an error correction model (ECM) used for the time series analysis, and Section 6 presents results from the ECM. Section 7 provides concluding remarks.

2. Quality of Life and Quality of the Business Environment

2.1 Theory: Workers and Firms

We adopt an open city model and focus initially on workers and firms, both of who are assumed to be mobile. In addition, we assume that all lots are identical, all workers are equally skilled, and all workers inelastically supply one unit of labor. A spatial equilibrium across cities requires that worker utility (u) and firm profit (π) both be equal across metropolitan areas (j = 1, …, J) as given by

$$\bar{u} = u(w_j, r_j \mid A_j) \tag{2.1}$$

and

$$\bar{\pi} = \pi(w_j, r_j \mid A_j)\,. \tag{2.2}$$

In (2.1) and (2.2), w_j is the wage in city j relative to a given reference city, the wage for which is normalized to 1. Similarly, r_j is the land rent in city j relative to the reference city, the land rent for which is also normalized to 1. The vector of amenities in city j is given by A_j, while $\bar{u}$ and $\bar{\pi}$ are the equilibrium levels of utility and profit in the system. A third equilibrium condition requires that city population levels (P_j) sum to total population, $\bar{P}$,

$$\bar{P} = \sum_j P_j\,. \tag{2.3}$$

Expressions (2.1) and (2.2) can be solved for the equilibrium wages and land rents in each city as shown in Figure 1 (see also Blomquist et al (1987) and Gyourko and Tracy (1991)). Holding A_j constant in city j, the iso-utility curve, $\bar{u}_j$, traces out the set of wages and land rents that satisfy (2.1) for city j: this function is upward sloping because higher w_j must be offset by higher r_j. Similarly, the iso-profit curve in city j, $\bar{\pi}_j$, traces out the set of w and r that satisfy (2.2): this function is downward sloping because higher *w* must be offset by lower r_j. The intersection of $\bar{u}_j$ and $\bar{\pi}_j$ yields w_j* and r_j* for all *j*, … J, the equilibrium wages and land rents in each city, while simultaneously ensuring that (2.3) is satisfied.

Next, totally differentiate the indirect utility function along an indifference curve as in Roback (1982) to get

$$dw_j u_w + dr_j u_r + dA_j u_A = 0 \ .$$

Rearranging and applying Roy's Identity yields

$$u_A / u_w = L^* \cdot dr_j / dA_j - dw_j / dA_j \ ,$$

where *L** is the optimum land use per household, while u_A / u_w is the marginal rate of substitution between *A* and *w*, or equivalently, the amount of income a household is willing to give up in exchange for a unit increase in *A* (with inelastic labor supply). Define $u_A / u_w \equiv b_{h\,j}$ to simplify notation and normalize *L** to 1. Then the vector of marginal benefits that households derive from A_j is

$$b_{h\,j} = dr_j^H / dA_j - dw_j / dA_j \, . \qquad (2.4)$$

where r^H is the rent on a normalized unit of land per household. Equation (2.4) is the same expression as in Roback (1982, 1988), Blomquist et al. (1988), and Gyourko and Tracy (1991).

Recall now that r_j and w_j vary across cities only because of differences in amenities. In addition, r_j and w_j are defined relative to a reference city, the land rent and wage rate for which are both normalized to 1. Then pre-multiplying both sides of (2.4) by A_j, the benefit that households derive from amenities in city *j* relative to the reference city can be written as

$$Q_{H_{jt}} = r_j^H - w_j \ . \qquad (2.5)$$

Observe that (2.5) describes the amount of real wage families would be willing to give up to live in city j, where the real wage decreases with r^H and increases with *w*. This measure is referred to as the urban "quality of life" in the literature.

A city's amenities affect not only household welfare, but also the cost of doing business. For example, firms trading with Canada will favor northern locations while those trading with Asia will favor west-coast locations, *ceteris paribus.* Similarly, proximity to natural resources, city size, and fiscal policies and services all influence the cost of producing and marketing products.

The value firms place on a city's basket of amenities can be derived in an analogous manner to that above. To begin, re-write the profit function in (2.2) separating total revenue and total cost as

$$\pi(w_j, r_j \mid A_j) = \mathrm{xq} - \mathrm{x}c(w_j, r_j \mid A_j) ,$$

where q is the product price, x is output, and $c(w_j, r_j \mid A_j)$ is the cost function. Totally differentiating the indirect profit function along an iso-profit curve yields

$$c_r\, dr_j + c_w\, dw_j + c_A\, dA_j = 0 \ .$$

Then, rearranging and applying Shepard's Lemma

$$-c_A/c_w = \frac{L_j^*}{N_j^*} \cdot dr_j / dA_j + dw_j / dA_j \ ,$$

where $-c_A/c_w$ is the ratio of the impact on production costs from a unit change in A relative to that of a unit change in labor, or equivalently, the additional input cost a firm is willing to incur in exchange for a unit increase in A. Note also, that L^*/N^* is the optimal ratio of land per worker. Next define $-c_A/c_w \equiv b_{f\ j}$ to simplify notation and normalize the amount of land per worker to 1. Then the vector of marginal benefits firms derive from A_j is given by

$$b_{f\ j} = dr_j^F / dA_j + dw_j / dA_j \ , \qquad (2.6)$$

where r^F is the normalized level of land use per worker.

Pre-multiplying both sides of (2.6) by A_j, the benefit that firms derive from amenities in city *j* relative to the reference city can be written as

$$Q_{F_{jt}} = r_{jt}^F + w_{jt} \ . \qquad (2.7)$$

Observe that (2.7) describes the additional input costs firms are willing to incur to locate in city j, where input costs increase with both r^F and w. To complete the analogy with households, hereafter Q_F is referred to as the urban "quality of business environment."[9]

The comparative statics of the model above with respect to the impact of amenities (A) on the equilibrium distribution of workers across cities is complicated because individual amenities can have similar or opposite appeal for households and firms.[10] Because an important goal of this paper is to evaluate the time series relationship between amenity valuations and net migration across cities, the model above is now reformulated by writing the household and firm equilibrium conditions, (2.1) and (2.2), as

$$\bar{u} = u(w_j, r_j \mid Q_{H_j}) \tag{2.8}$$

$$\bar{\pi} = \pi(w_j, r_j \mid Q_{F_j}), \tag{2.9}$$

where $\partial u/\partial Q_H$ and $\partial \pi/\partial Q_F$ are both positive. In both (2.8) and (2.9), a change in the amenity vector itself *or* a change in tastes both serve to affect the value that households and firms place on a city's amenities. Moreover, Q_H and Q_F both have unambiguously positive effects on household utility and firm profits, and therefore, positive effects on a city's population of workers.

2.2 Theory: Retirees

The model above can be extended to examine the influence of amenities on retiree locations. A subtle issue arises, however, with regard to how to measure the value that retirees place on a metropolitan area's basket of amenities. Because retirees do not participate in labor markets, it is tempting to ask whether inter-city differences in land rents alone can be used to measure retiree preferences for different

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[9]From (2.5) and (2.7) it is clear that whereas household valuations of amenities are given by the *difference* between r and w, firm valuations of amenities are given by the *sum* of r and w. That difference arises because the iso-utility and iso-profit functions discussed above have opposite slopes.

[10]See the Appendix to Blomquist et al (1988), for example.

locations. For two reasons this is not likely to be the case. First, for the cities included in our sample, retirees are a small part of the population relative to working households. For that reason, our model is predicated on the assumption that local amenity effects are capitalized into both equilibrium wages and land rents through their influence on workers and firms in the manner portrayed in Figure 1. Second, retirees face the same equilibrium land rents (and wages) as all other members of the population.

To understand how these features affect measurement of retiree amenity valuations, consider the following example. Suppose first that retirees were the dominant players on the demand side of the land market. In that case, land rents would always be higher in areas favored by retirees because of outward shifts in the retiree demand for land, and retiree amenity valuations would not depend on wage rates. But, when workers and firms dominate the demand side of the land market, such a monotonic correspondence between retiree amenity valuations and land rents would not prevail. For example, from Figure 1, if a city is attractive to both households and firms then land rent will be high. But if the city is attractive to households and unattractive to firms then land rents could be low. Because the preferences of firms affect both land rents and wages, land rents only partially capture the value of local amenities, the remaining portion of which is reflected in the market wage. Measures of retiree amenity preferences based on market equilibrium land rents, therefore, must also take equilibrium wages into account as well.

In light of these arguments, and to facilitate the analysis to come, hereafter we assume that retirees place the same value on a city's basket of amenities as do workers, both of which are measured by Q_H as defined in expression (2.5). Retiree indirect utility is then represented as

$$\bar{u}_R = u(r_j, Q_{H_j}) \ , \qquad (2.10a)$$

where $\partial u(r_j, Q_{H_j})/\partial r < 0$ because an increase in rent reduces retiree purchasing power, while $\partial u(r_j, Q_{H_j})/\partial Q_H > 0$ for reasons outlined above, and $\bar{u}_R$ is the equilibrium level of utility among retirees. Observe also, that the *w* does not appear in (2.10a) because retirees do not participate in the labor market.

Recall next that r_j is determined by Q_H and Q_F as shown in Figure 1. Substituting for *r* in (2.10a), the retiree's indirect utility function can also be written as

$$\overline{u}_R = u(Q_{F_j}, Q_{H_j}) \ . \qquad (2.10b)$$

As a city becomes more attractive to industry, labor demand shifts out pushing land rent up. This hurts retirees and hence, $\partial u(Q_{F_j}, Q_{H_j}) / \partial Q_F < 0$. Analogously, as Q_H increases a city becomes more attractive to households, but this also causes labor supply to shift out pushing land rent up. As a result, the sign of $\partial u(Q_{F_j}, Q_{H_j}) / \partial Q_H$ is indeterminate.

To summarize, in the model described by Figure 1 and expressions (2.1) to (2.10), workers and firms migrate in response to inter-city differences in Q_H and Q_F. That migration (or threat of migration), determines the equilibrium set of land rents and wages across cities while simultaneously determining the equilibrium distribution of workers *and* retirees across metropolitan areas. The distribution of workers is governed directly by shifts in the labor supply and demand curves for individual cities. The distribution of retirees is *indirectly* influenced by shifts in the labor supply and demand curves because retirees, workers, and firms all compete for space.

2.4 Empirical Measures

Estimates of the amenity valuations are constructed as follows. As in Blomquist et al (1988) and Gyourko and Tracy (1991), wage and building rent hedonic regressions are specified as[11]

$$\log(w_{ij_t}) = \alpha_{w_{ot}} + \alpha_{w_{1t}} Z_{ij_t} + \gamma_{w_{j_t}} D_{ij_t} + u_{w_{ij_t}} \qquad (2.11)$$

and

[11]Although the model established earlier in this section is specified in terms of land rents, as shown in the literature, utilizing building rents as below does not affect the structure of the analysis or the interpretation of the results (see Blomquist et al. (1988), for example).

$$\log(r_{ij_t}) = \alpha_{r_{ot}} + \alpha_{r_{1t}} X_{ij_t} + \gamma_{r_{j_t}} D_{ij_t} + u_{r_{ij_t}} \tag{2.12}$$

where Z_{ijt} is a vector of worker traits for individual *i* in city *j* and year *t*, and X_{ijt} is a vector of structural attributes of the building. Previous quality of life studies augment these variables with a long list of city specific amenities in order to represent a city's attributes. That approach, however, suffers from potential omitted variable bias because important amenities may be left out.[12] In addition, the standard approach presents large data requirements given that values for the time-varying amenities must be obtained for each time period. As an alternative, we include a vector of metropolitan area fixed effects, D_{jt}. Having controlled for the quality of the worker's skill level and the building's structural attributes through Z and X, the estimated fixed effects (γ_{wjt} and γ_{rjt}) reflect all location-specific attributes that affect intermetropolitan spatial variation in wages and property values at time *t*.[13]

Expressions (2.11) and (2.12) are estimated separately for each time period in the sample. This yields a panel of estimated fixed effects, γ_{wjt} and γ_{rjt}, which are used to form Q_H and Q_F. This is done by first taking derivatives of the anti-logs of (2.11) and (2.12) with respect to location (D_j)

$$w_{j_t} \equiv \frac{\partial w_{j_t}}{\partial D_{j_t}} = \hat{\gamma}_{w_{jt}}\, e^{[\,\hat{\alpha}_{w_{0t}} + \hat{\alpha}_{w_{1t}} Z_{j_t} + \hat{\gamma}_{w_{j_t}} D_{j_t}\,]} \tag{2.13}$$

and

$$r_{j_t} \equiv \frac{\partial r_{j_t}}{\partial D_{j_t}} = \hat{\gamma}_{r_{j_t}}\, e^{[\,\hat{\alpha}_{r_{0t}} + \hat{\alpha}_{r_{1t}} X_{j_t} + \hat{\gamma}_{r_{j_t}} D_{j_t}\,]}\,. \tag{2.14}$$

Next, Z, X, and D are fixed at reference values such that the only variation in w_{jt} and r_{jt} over time and across cities is through the estimated panel of $\hat{\gamma}_{w_{j_t}}$ and $\hat{\gamma}_{r_{j_t}}$, consistent with the theory above.

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[12]Gyourko and Tracy (1991), for example, emphasize that Blomquist et al (1988) erred by not including local fiscal amenities in their hedonic regressions. When Gyourko and Tracy (1991) include local fiscal amenities in their model, they find that fiscal amenities are at least as valuable to households as non-fiscal amenities. That result suggests that local government can enact policies that affect Q_H.

[13]In (2.11) and (2.12) we assume that the coefficients on *Z* and *X* are similar across cities but vary over time.

Substituting (2.13) and (2.14) evaluated at the reference values into (2.5) and (2.7) gives the quality of life and quality of business environment, $Q_{H,jt}$ and $Q_{F,jt}$, for each city and period.

3. Data

Data required for the estimation of equations (2.13) and (2.14) were obtained from the March outgoing rotation files of the Current Population Survey and the national core files of the American Housing Survey, respectively, over the 1977 to 1995 period. From those data, Z, X, and D in (2.13) and (2.14) were set equal to their sample means for 1980, the 1980 average annual wage earnings and building rents in the population. That is the same reference base as used by Blomquist et al. (1988) and Gyourko and Tracy (1991) and facilitates comparisons to their work.

While the CPS data were obtained annually for each year from 1977 to 1995, the AHS data were available on an annual basis only for the years from 1977 to 1983. After 1983, Census collected the AHS data on a biannual basis. To fill in the missing years, quality adjusted building rents for 1984, 1986, 1988, 1990, 1992, and 1994 were linearly interpolated from the adjacent years.

The wage variable employed in the study is readily obtained as the total annual wage and salary earnings for the worker. In contrast, construction of a rent variable is more complicated. Annual information on r^F is not available while r^H can be readily observed in the AHS. For that reason, estimates of r^H and r^F are both based on rents for residential properties, effectively setting r^H equal to r^F. In estimating U.S. GDP, the Commerce Department similarly is constrained by the absence of a quality-adjusted non-residential real estate price series. Accordingly, our approximation here is consistent with the Commerce Department practice of using quality-adjusted residential price indexes to estimate the price deflators for both residential and non-residential real estate. Following that practice, we calculate gross rents based on householder gross rents for renter-occupied units and owners’ estimates of house value for owner-occupied units. In the later case, as with Gyourko and Tracy (1991) and Blomquist et al (1988), owners’ estimates of property value are converted to annual rents using a discount rate of 7.85 percent taken from Peiser and Smith's (1985) paper on owner-occupied housing user costs.

To be included in the wage sample, an individual needed to be a full-time and full-year worker earning in excess of $1,000 per year. When estimating the rent hedonic, excluded from the housing sample were mobile homes, public housing units, rent controlled units, and other government subsidized units. In both cases, to be included in the sample an observation (individual or housing unit) had to be located in an identified MSA. The resulting sample sizes of the housing expenditure and wage samples used for the analysis vary somewhat across both the CPS and AHS data sets and the years of analysis. In 1978, for instance, the housing expenditure and wage analyses were based on samples of 23,734 and 13,981 observations, respectively.

Worker earnings were quality-adjusted using information on the worker's educational achievement, age, race, marital status, and number of dependents. Housing expenditures were quality-adjusted in a similarly standard manner using information from the AHS on the housing unit characteristics and characteristics of the neighborhood.[14] In general, results from the wage and rent hedonic regressions conform to well-established findings in the literature.[15]

Population data for different metropolitan areas were obtained from a variety of Census Department publications but primarily from the State and Metropolitan Area Data Books and various issues of the Statistical Abstract of the United States. The data were collected on a county-level basis and

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[14]Wage hedonic regressors included age and age squared of the individual, age and age squared of the spouse, white versus non-white, number of children under aged 6 in the family, and number of children between age 6 to 18 in the family. In addition, each of the age variables for both the individual and spouse were separately interacted with four education categories: high school degree, some college, 4-year college degree, and more than a college degree, where less than a high school was the omitted category. Housing hedonic regressors included whether the unit was single family detached, attached or multi-family, number of rooms, number of bedrooms, presence of a garage, presence of a basement, number of bathrooms, central air conditioning, room air conditioning, central heat, abandoned buildings on the street, age of building, whether HUD characterizes the building as being in a dilapidated condition, and central city versus non-central city location.

[15]In total, 38 hedonic regressions were run. Results from those regressions are not presented to conserve space.

aggregated to get the metropolitan area population levels using the 1993 Census definitions of the metropolitan areas. After cleaning both the population data and the samples used to construct the quality of life and quality of business data, we were able to obtain a complete panel of annual observations for thirty-seven cities from 1977 to 1995.[16]

4. **An Initial View of the Quality of Life and Quality of Business Environment Series**

Table 1 displays quality of life and quality of business environment measures (Q_H and Q_F) for the 37 cities for which a complete set of data could be constructed. Average values are based on biannual observations to reduce spurious correlation when calculating the standard errors for the Q_H and Q_F values when averaging across years[17] All values are in 1980 dollars and are constructed as previously described. All columns are sorted in ascending order based on the 1977-95 average values of Q_H.

Focus first on the 1977 to 1995 average values for Q_H. Observe that the range in estimates from lowest to highest is roughly $7,500 while the interquartile range (from 25th to 75th percentile) is roughly $2,000. These values are close to those of Gyourko and Tracy (1991) who report comparable numbers of $8,227 and $1,484 based on their cross-sectional analysis of 1980 census data. In addition, casual comparison of our 1977-95 average values and rankings and those of Gyourko and Tracy (1991) suggest that there is a relatively high degree of correlation in findings. In both studies, older, industrialized

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[16]In the 1970s, the CPS identified only the 39 largest cities in the United States. In two of those cities it is not possible to obtain measures of population within a fixed set of geographic boundaries over time. This reduced our set of cities to the 37 noted above.

[17]Standard errors for the Q_H and Q_F values in Table 1a were calculated based on the estimated covariances for the hedonic coefficient from the different years over which the hedonic models were estimated. In addition, the average values for Q_H and Q_F in Table 1a reflect values taken from every other year of the Q_H and Q_F measures beginning with 1977. This was done because the CPS sample turns over every two years. In addition, most renters move within two years of arriving in their home, although most homeowners do not. As an approximation, therefore, biannual data largely removes spurious correlation when averaging Q_H and Q_F over time. That permits us to set the covariance terms to zero when calculating the variance and standard errors for the average values of Q_H and Q_F reported in Table 1a.

midwestern and eastern cities such as Detroit, Kansas City, Baltimore, Cleveland, Cincinnati, Gary, and Akron, were ranked among the lowest quality-of-life metropolitan areas. In addition, in both studies warm coastal cities like Miami, San Diego, San Francisco, San Jose, Sacramento, and Los Angeles-Anaheim score among the areas with the highest quality of life. Thus, it appears that our 1977-95 average Q_H measures are relatively similar to those of Gyourko and Tracy (1991), despite the use of different data and a different estimation method.[18]

A striking result now emerges when one compares household and firm valuations of different cities. Observe that many of the cities least attractive to households are large industrial metropolitan areas, such as Detroit, Gary, Baltimore, Philadelphia, and Washington D.C, all of which are ranked in the bottom quartile by households. How could these cities be among the largest metropolitan areas in the United States if households don't want to live there? The answer of course, is that firms value the amenities in these areas: all of these cities are ranked in the upper half by firms. Detroit, for example, was ranked 37 by households but was ranked 9 by firms. Washington D.C was ranked 25 by households but 4 by firms. Two important messages emerge from this comparison. First, for a city to grow large, either households must want to live in the city as in Miami – pushing labor supply out – or firms must want to do business in the city as in Washington D.C. – pushing labor demand out – or both as in New York, San Francisco, and Los Angeles. Second, firms and households often do not like the same cities, a result that reflects the very different objectives of the two groups – utility maximization versus profit maximization.

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[18]The comparisons above are based on the second model presented in Gyourko and Tracy (1991): "Random Effects, Group Effects Included" in Table 3 of their paper. That model is the closest to the fixed effects approach used here. In addition, the median standard error of Q_H and Q_F across cities and years in our sample was roughly $1,200 (because Q_H and Q_F are estimated as linear combinations of r_j and w_j, the standard errors on Q_F are identical to those for Q_H). That estimate is also close to standard errors for the most comparable model presented in Gyourko and Tracy (1991). In contrast, the median standard error for the 1977-95 biannual average for Q_H and Q_F in Table 1a is considerably smaller, $342. The greatly reduced standard errors in Table 1a reflect the greater quantity of data used to calculate the biannual average values relative to those for individual cities and years.

Do these patterns persist over time? To address that question, in Table 1b we regress the biannual average for Q_H over the 1987 to 1995 period on the biannual average from the 1977 to 1985 period. An analogous regression is also reported for Q_F. Observe that in the Q_H regression the coefficient on Q_H-Avg_{77-85} equals .866 with a t-ratio of 11.97, while the R^2 is .80. Similar values are obtained for the Q_F regression: the coefficient on Q_F-Avg_{77-85} equals 1.07 while the R^2 is .72. These findings indicate that cities that were relatively attractive (unattractive) to households and firms during the 1977-85 period tend to remain relatively attractive (unattractive) in the 1987-95 period. By the same token, changes in quality of life and quality of the business environment do occur. This is apparent from the plots of Q_H and Q_F in Figures 2 and 3. To what extent then do changes in Q_H and Q_F influence the relative size and composition of cities in the economy as implied by the model in Section 2? It is to this question that we now turn our attention.

5. **Empirical Model: Changes in Amenity Valuations and Migration in a System of Cities**

5.1 The error correction model

The principal framework used to analyze the system of cities model described in Section 2 is an error correction model (ECM). This model takes advantage of the year-to-year variation in the data and is comprised of two parts. The first part, typically referred to as the cointegrating equation, describes the long run relationship between the variables of interest, in this case, city population shares, Q_H, and Q_F as outlined in Section 2. The second part, referred to as the error correction equation, is a vector autoregression (VAR) in which the variables of interest are constrained to return to their long-run relationship following a shock to the system. This equation permits us to estimate the speed of such convergence. The development of these equations proceeds as follows.

Consider first plots of worker log-population shares in Figure 4 for the individual cities in the sample. Observe that most of the cities display rather pronounced trends. A similar pattern holds for retiree log-population shares in Figure 5. These plots suggest that it is desirable to control for trends when evaluating changes in population shares over time.

Next, augmented Dickey-Fuller (ADF) tests were estimated to check for unit roots in each of the principle series to be studied. In Table 2 observe that one cannot reject the null hypothesis that worker and retiree log-population shares, Q_H, and Q_F are I(1), and hence, are non-stationary and non-mean reverting series.[19] That result is important because if log-population shares were of a different order of integration – the degree of differencing necessary to render the series stationary – then Q_H and Q_F could not explain intertemporal flows of population across cities.[20] Finding that each of the principal series are of the same order of integration, therefore, is a necessary condition if Q_H and Q_F are to determine the spatial distribution of population across cities.[21] In light of these arguments and results, we treat all of the principal series as I(1). As will become apparent, this also has implications for the specification of the ECM.[22]

We turn now to the cointegrating equation that governs the long-run relationship between worker log-population shares (P^{worker}), Q_H, and Q_F , variables that are expected to move together over time to equilibrate the system of cities. As will become apparent, analogous regressions are also run for retiree log-

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[19]Each ADF test was conducted with a constant to control for drift and a time trend. Two lags of the dependent variable were also included to control for serial correlation. Diagnostic tests of the residuals from the unit root regressions indicate that these controls were sufficient to whiten the residuals, a necessary condition for valid interpretation of the ADF test statistics. Jarque-Bera tests of the residuals from the ADF regressions could not reject the null of a normal distribution in all but a couple of cases, while Q-statistics indicated flat correlograms in nearly all instances. In addition, the trend term was often highly significant, and omitting the trend term often caused the residuals to have a pronounced trend, especially in the population share regressions.

[20]Regressing an I(1) series on an I(0) series, for example, always yields a zero coefficient with a sufficiently long time series. The reason is that the I(1) series follows a random walk and is not mean reverting, while the I(0) series is stationary and always reverts back to its long-run mean given a sufficiently long time series.

[21]In addition, evidence that log(P^{worker}) is I(1) is exactly what one would anticipate in a system of cities where larger cities gain a comparative advantage over competing metropolitan areas because of enhanced urbanization economies, and therefore, grow larger still. Helsley and Strange (1994), for example, argue that even if cities have similar initial endowments, those cities that develop first – encounter positive initial shocks to their population shares – ultimately become the largest cities in the system because of the agglomeration economies that stem from the early manufactured infrastructure.

[22]As a robustness check, in Table A-1 of the Appendix we also present ADF tests on the log-population shares for six cities based on decade-by-decade data going back to 1800. Although the longer time horizon increases the power of the ADF tests, we still fail to reject the null of a unit root in five of the six cities.

population shares by substituting $P^{retiree}$ for P^{worker}. After some experimentation, our preferred specification of the cointegrating relation is given by the following equation,

$$\log(P_{jt}^{worker}) = a_{o_j} + c_1 Q_{H_{jt}} + c_2 Q_{F_{jt}} + a_{1_j}\Delta Q_{H_{j+1,t}} + a_{2j}\Delta Q_{H_{j-1,t}} + a_{3j}\Delta Q_{F_{j+1,t}} + a_{4j}\Delta Q_{H_{j-1,t}} + a_{5j} t + e_{p,jt}, \quad (5.1)$$

for all cities $j = \{1, \ldots, J\}$. The coefficients of interest are c_1 and c_2. Evidence that these coefficients are both positive would imply that amenity-induced outward shifts in the labor supply and demand curves serve to increase the equilibrium size of a city's workforce, a central prediction of the model in Section 2 and Figure 1. Alternatively, when $P^{retiree}$ is substituted for P^{worker} we expect c_2 to be negative while the sign of c_1 is indeterminate for reasons outlined earlier.

To ensure that (5.1) yields consistent estimates of c_1 and c_2, a city-specific constant is included in (5.1) to control for time-invariant city effects, while a time trend is included given evidence of trending as noted above.[23] In addition, because the principal series are all assumed to be I(1), OLS estimates of the coefficients on Q_H and Q_F would ordinarily have non-standard limiting distributions making inference difficult [e.g. Engle and Granger (1987)]. To address that problem, (5.1) includes leads and lags of the first-differences of the coefficients on Q_H and Q_F. As shown by Saikkonen (1991), this causes the standard errors on c_1 and c_2 to have standard limiting distributions allowing us to interpret the OLS t-ratios in the usual way. Finally, when estimating (5.1), c_1 and c_2 were constrained to be alike across cities in order to increase the power of our tests given that we only have 19 time periods in the sample.[24] In contrast, all

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[23]Note also, that if measurement error in Q_H and Q_F can be approximated as a city-specific effect, then inclusion of city fixed effects greatly reduces the influence of measurement error that might otherwise bias estimates of c_1 and c_2 towards zero.

[24]As shown by Stock (1987), if population shares, Q_H and Q_F are I(1) and are cointegrated, coefficient estimates of c_1 and c_2 converge at rate T instead of the usual square root of T which adds to the power of our estimates.

other variable coefficients in the model were allowed to vary across cities in order to allow for differences across metropolitan areas.[25]

The vector error correction equation corresponding to (5.1) is given by,

$$\Delta \log(P_{jt}^{worker}) = \theta_{o_j} + ge_{j,t-1} + \sum_{l=1}^{2} \theta_{P_j} \Delta \log(P_{j,t-l}^{worker}) + \sum_{l=1}^{2} \theta_{Q_{H\,j}} \Delta Q_{H_{j,t-l}} + \sum_{l=1}^{2} \theta_{Q_{F\,j}} \Delta Q_{F_{j,t-l}} + \varepsilon_{jt}, \tag{5.2}$$

where $\Delta \log(P_{jt}^{worker})$ is the annual growth rate of city j's worker population share and j = {1, …, J} (and $\Delta \log(P_{jt}^{retiree})$ is substituted into (5.2) when focusing on retirees). Two lags of each of the endogenous variables are included in (5.2) to control for serial correlation arising from the short run dynamics. A constant is also included since a time trend is included in the cointegrating equation. Coefficients on these variables are allowed to vary across cities to allow for differences across urban areas as before.

A key variable in (5.2) is the 1-period lag of the error term from (5.1), $e_{j,t-1}$.[26] Given that (5.1) describes the long run relationship between P^{worker}, Q_H and Q_F, the lagged error term can be interpreted as a short run deviation of the system cities from its long-run equilibrium path. Provided $e_{j,t-1}$ has a non-zero coefficient in (5.2), its inclusion in the model constrains $\Delta\log(P_{jt})$, the dependent variable in (5.2), to return to its long run path. Moreover, the coefficient on $e_{j,t-1}$, denoted by g, measures the speed of convergence back to long run equilibrium. Because each city j = {1, …, J} is part of the system, in general no city can

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[25]As discussed by Engle and Granger (1987), log-population shares, Q_H, and Q_F are cointegrated if there exists at least one linear combination of the variables that yields a stationary outcome. That property can be tested by applying ADF tests to the residuals from (5.1). Results of those tests are presented in Table A-2 of the Appendix. In 27 of the 37 cities we reject the null of a unit root in favor of stationarity. Similar tests produce comparable results for alternative specifications of (5.1) to follow. These results support the presence of a stable long-run relationship between the variables of interest for most but not all cities.

[26]Including $e_{j,t-1}$ in (5.2) presumes that population shares and city valuations are cointegrated. Otherwise, one could omit $e_{j,t-1}$ and estimate (5.2) based on an unrestricted VAR. As will become apparent in Table 3, however, the coefficient on $e_{j,t-1}$ is highly significant which argues for inclusion of $e_{j,t-1}$ in the model. In addition, as noted previously, ADF tests presented in Table A-2 of the Appendix reject the null that $e_{j,t-1}$follows a unit root for most cities, consistent with the assumption that population shares and city valuations are cointegrated.

be in long run equilibrium unless the entire system is in equilibrium. For that reason, when estimating (5.2) *g* is constrained to be alike for all cities.[27]

6. Results from the ECM for Workers and Retirees

Table 3 provides the core results for the ECM.[28] In column 1, worker log-population shares are regressed on Q_H and Q_F based on the specification in (5.1). Observe that both Q_H and Q_F have strong positive and highly significant effects on a metropolitan area's share of the workforce. This supports the theory underlying Figure 1: as a city becomes relatively more attractive to workers and firms, labor supply and demand both shift out and the relative size of the city in the overall system of cities increases.

In column 2 retiree log-population share is substituted for worker log-population share. The specification of the model is otherwise the same as before. In sharp contrast to column 1, Q_H and Q_F have negative and significant effects on retiree population shares, consistent with the theory underlying equation (2.10b). To further explore this result, the model is re-estimated in column 3 substituting land rent, *r*, for Q_F. Results from this model indicate that retiree population shares fall with an increase in land rent, while Q_H has a positive but insignificant coefficient. Apart from the insignificance of the coefficient on Q_H, these findings are consistent with the theory underlying equation (2.10a).

These latter results are reinforced by the last two columns of the table where the log ratio of retiree to worker population shares is used as the dependent variable. That ratio is equivalent to $\log(R_j / N_j) - k$,

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[27]Together, two-stage estimates of equations (5.1) and (5.2) comprise a modified Engle-Granger (1987) method for estimating a vector error correction model. An alternative to the Engle-Granger method is the Johansen method (Johansen (1991)) which is a full information Maximum-Likelihood procedure capable of estimating both the ECM and the number of cointegrating vectors. However, the Johansen method suffers from substantial small sample bias and lack of robustness to specification error. Although the Engle-Granger method requires that one impose *a priori* the number of cointegrating vectors (of which there could be either one or two in our case), the Engle-Granger method is more robust in the presence of short time series such as ours.

the ratio of retirees to workers in city j less a constant *k*, where *k* equals the log-ratio of retirees to workers for the entire system of cities (denoted by $\log(\sum_j R_j / \sum_j N_j)$). Cities for which $\log(R_j / N_j) - k$ is high, therefore, have a high ratio of retirees to workers relative to the overall system.

In column 5 observe that Q_H and Q_F have strong negative and highly significant effects on the ratio of retirees to workers in a metropolitan area. In column 6, as before, an increase in the cost of living (*r*) has a negative and significant impact on the relative presence of retirees in a city as compared to workers. In addition, an increase in Q_H has a positive and significant influence on a city's retiree-to-worker ratio. These findings lend further support to the argument that retirees tend to seek out attractive low-cost locations, consistent with theoretical arguments by Graves and Knapp (1988). More generally, the pattern of results suggests that retirees tend to move away from cities that become increasingly attractive to workers and firms in order to avoid the higher land rents that result from increased competition for space.

Table 3 also presents estimates of the coefficient on the error correction term in the second stage equation as described in equation (5.2). For the worker population models and the retiree/worker population models, the coefficients are roughly .2. This says that following a shock to the system of cities that creates migratory pressures – as with an influx of immigrants to a given city or with the enactment of new pro-business policies, for example – the distribution of population in the system moves roughly 20 percent of the way back to long run equilibrium after 1 year. This translates into an adjustment period of roughly 8-1/2 years for 90 percent return to equilibirum. In contrast, the error correction coefficient for the retiree models in columns 2 and 3 is roughly .3. This implies a period of roughly 6 years for 90 percent adjustment back to equilibrium. The clear message here is that retiree population shares respond to unexpected shocks to the system of cities roughly one-third faster than the population of workers. Given

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[28]To conserve space, only estimates of c_1, c_2, and *g*, are presented, the coefficients of primary interest. The remaining coefficients in the model all vary across cities and total 444 parameters over the 37 cities for equations (5.1) and (5.2).

that migration of workers requires a coordinated movement of industry-specific human and physical capital, the slower adjustment period for workers is quite plausible.

7. Conclusion

Are the cities in which families want to live similar to the cities in which firms wants to operate? To what extent do changes in the value of local amenities affect the relative size and composition of individual cities, especially with regard to the presence of workers and retirees? This paper addresses these and related questions while also making three important methodological contributions to the literature on urban quality of life.

Our analysis is based on the argument that urban amenities are vitally important to the locational choices of both households and firms. However, while existing literature has measured household valuations of urban areas – referred to as quality of life in the literature – systematic assessment of the firm valuations of urban areas – referred to as quality of the business environment in this paper – has been absent from the academic literature. Accordingly, our most important methodological contribution is to construct and estimate a measure of firm valuations of city-specific amenities in a manner that is solidly grounded in economic theory.

A second important innovation concerns dynamics. Prior to this paper, the quality of life literature has been largely static in nature, precluding analysis of changes in city valuations over time and related effects on the growth and composition of individual metropolitan areas. In this paper, we construct a panel of household and firm valuations of urban amenities for 37 metropolitan areas in the United States from 1977 to 1995. That panel is developed using panel data methods that greatly simplify construction of amenity valuations relative to prior studies. This allows us to examine the dynamics of worker and retiree migration across cities in response to changes over time in the quality of life and the quality of the business environment.

Our third important innovation concerns tests of the model. Prior studies have emphasized that individual amenities could have opposite or similar appeal to workers and firms, and therefore, ambiguous effects on the equilibrium distribution of population across cities. That emphasis has largely precluded testing of the model. In contrast, by focusing on the relationship between population shares across cities, quality of the business environment, and quality of life, we are able to test several of the most fundamental predictions of the model.

A striking result is that many of the cities least attractive to households are attractive to firms. Detroit, Gary, Baltimore, Philadelphia, and Washington D.C, for example, are all ranked in the bottom quartile in the eyes of households but are all ranked in the upper half by firms. Those differences reflect the very different goals of the two groups – utility maximization versus profit maximization.

In addition, estimates from an error correction model (ECM) indicate that improvements in quality of life and quality of the business environment have strong positive effects on urban growth as measured by equilibrium city shares of workers, but negative effects on equilibrium city shares of retirees. The former result supports a central feature of the model, that worker population shares increase in response to amenity-induced outward shifts of the labor supply and demand curves. The latter result suggests that retirees avoid high land rents that result from in-movement of firms and workers. That finding is consistent with theoretical work by Graves and Knapp (1988) who argue that retirees should seek out cities in which amenities are capitalized into lower wages as opposed to higher land rents.

Additional results from the ECM indicate that following a shock that creates migratory pressures in the system of cities, worker-population shares converge 90 percent of the way back to long run equilibrium in 8-1/2 years while retiree-population shares achieve 90 percent convergence in 6 years. The longer response time of the worker population likely reflects the cost of moving industry-specific human and physical capital, a cost that impedes worker mobility relative to that of retirees.

Finally, from a policy perspective, results from this paper suggest an intriguing possibility. Economic development initiatives designed to attract workers and firms – such as tax breaks offered to industry – may have the unintended effect of repelling retirees. The degree to which this may or may not occur requires further research. Nevertheless, it is clear that local government policies designed to attract industry and retirees should ideally be coordinated so that one set of policy initiatives does not inadvertently negate the other.

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FIGURE 1: Equilibrium In the Land and Labor Markets

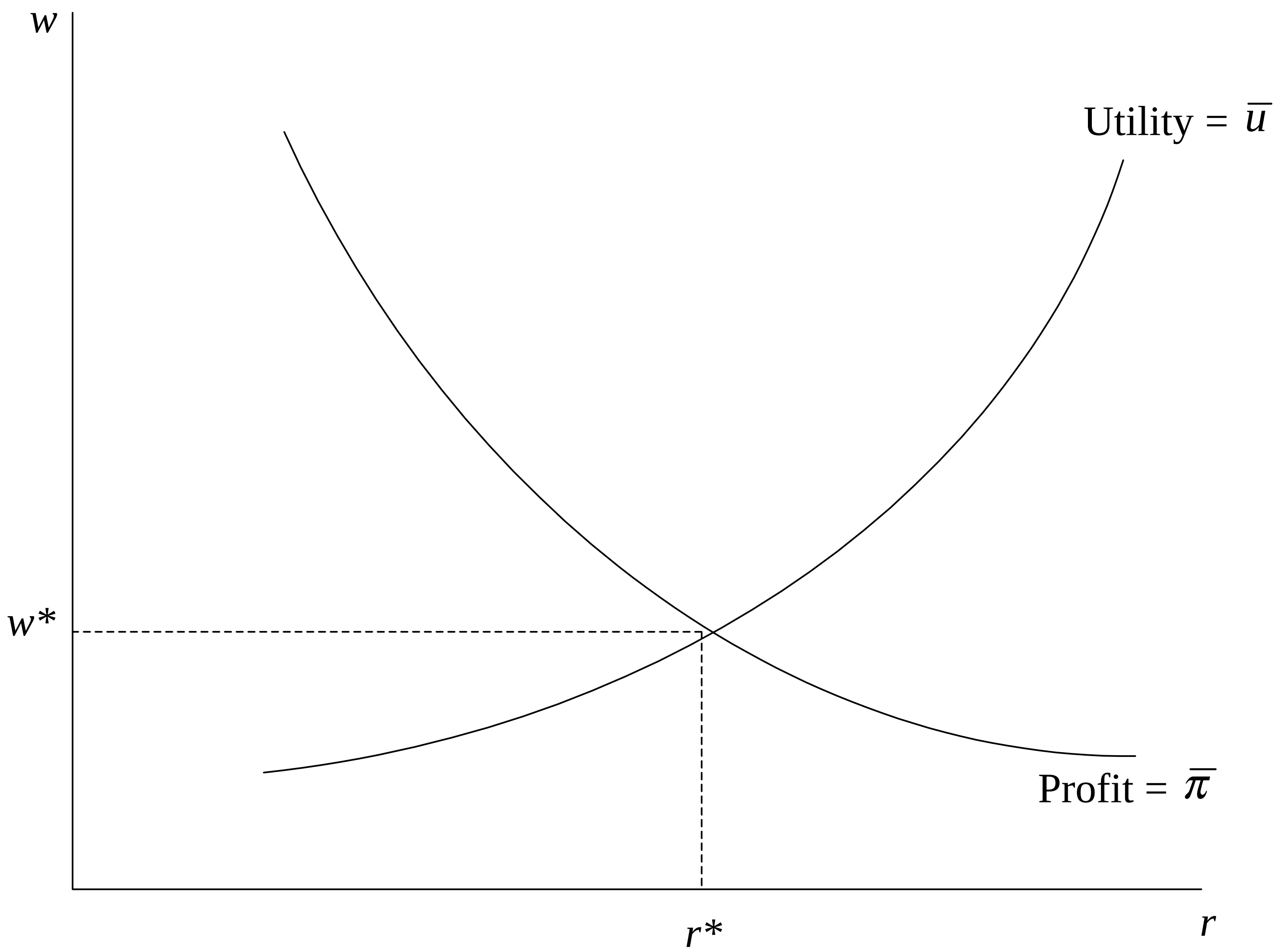

Table 1a
1977 to 1995 Average Values of Quality of Life and Quality of Business Environment*
(All values are in 1980 $; Rank = 1 is *best*, Rank = 37 is *worst*)

Differences in Q_H reflect the amount a household values one city over the other. Differences in Q_F reflect the amount a firm values one city over the other per worker

	Quality of Life (Q_H)			Quality of Business Environment (Q_F)		
Metropolitan Area	Rank Avg 77-95	Avg 77-95	Stnd Err	Rank Avg 77-95	Avg 77-95	Stnd Err
Miami	1	3632	327	34	-2111	327
San Diego	2	2385	346	10	1614	346
Los Angeles-Long Beach	3	2205	292	5	2710	292
San Francisco	4	2009	342	2	4786	342
Tampa-St. Petersburg-Clearwater	5	1728	339	37	-3202	339
New York	6	1606	292	7	2337	292
Albany-Schenectady-Troy	7	812	408	28	-1071	408
Greensboro--Winston-Salem--High Point	8	708	369	35	-2195	369
Sacramento	9	568	386	18	383	386
Norfolk-Virginia Beach-Newport News	10	312	393	30	-1158	393
Seattle-Bellevue-Everett	11	-3	346	6	2339	346
Denver	12	-52	331	15	807	331
Newark	13	-64	326	3	3791	326
San Jose	14	-274	361	1	5994	361
Minneapolis-St. Paul	15	-436	328	12	1246	328
Fort Worth-Arlington	16	-478	358	31	-1432	358
Birmingham	17	-504	392	36	-2786	392
New Orleans	18	-554	383	25	-524	383
Chicago	19	-658	294	8	1817	294
Indianapolis	20	-718	398	33	-1595	398
Rochester	21	-724	377	16	659	377
Pittsburgh	22	-781	333	29	-1075	333
Dallas	23	-797	322	20	52	322
Columbus	24	-813	355	26	-725	355
Washington DC	25	-871	298	4	3445	298
Milwaukee-Waukesha	26	-1111	355	14	845	355
Philadelphia	27	-1123	302	13	1168	302
Baltimore	28	-1145	336	11	1426	336
Cincinnati	29	-1247	345	23	-364	345
Atlanta	30	-1266	332	19	89	332
Cleveland-Lorain-Elyria	31	-1271	332	21	41	332
Akron	32	-1331	416	27	-851	416
Kansas City	33	-1389	338	32	-1578	338
Houston	34	-1401	313	22	-296	313
St. Louis	35	-1872	352	24	-427	352
Gary	36	-2719	533	17	548	533
Detroit	37	-3904	305	9	1657	305

*The *Q* averages were formed using every other year of the data beginning in 1977 to reduce spurious correlation when calculating the standard errors as discussed in the text.

Table 1b
Persistence of Amenity Valuations

	1987 to 1995 Biannual Average Q_H		1987 to 1995 Biannual Average Q_F	
	Coefficient	**t-ratio**	**Coefficient**	**t-ratio**
Constant	-429.1	-3.930	1,188.1	5.624
1977 to 1985 Biannual Average Q_H	0.8660	11.974	-	-
1977 to 1985 Biannual Average Q_F	-	-	1.0704	9.467
Observations	37		37	
R-squared	0.80		0.72	
Adj. R-squared	0.80		0.71	
Root MSE	662.59		1,281.0	

Figure 2: Quality of Life (2-Year Moving Average) – Q_H
(Vertical Scales Correspond to the Closest City in the Legend and Differ Across Plots)

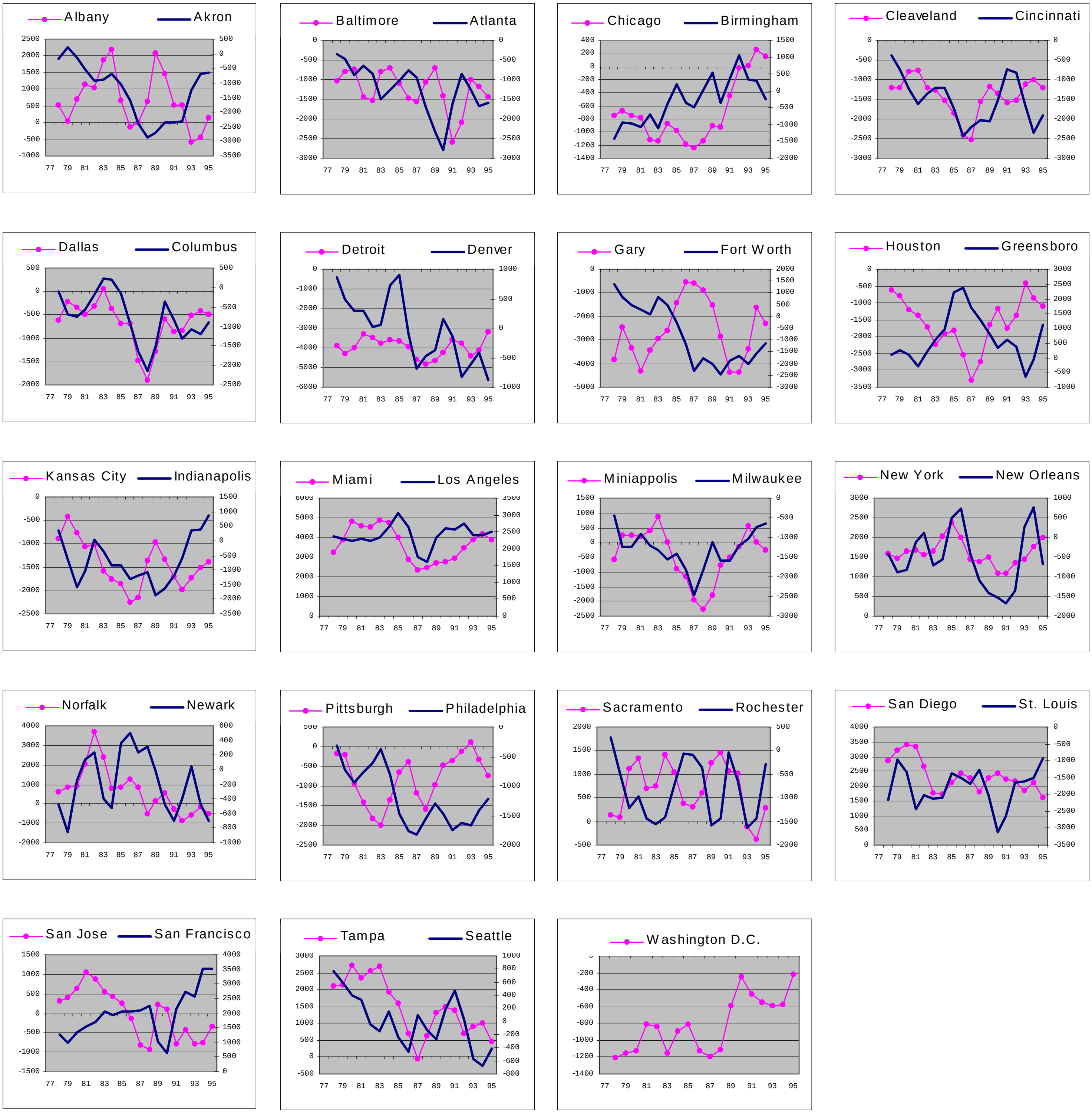

Figure 3: Quality of Business Environment (2-Year Moving Average) – Q_F
(Vertical Scales Correspond to the Closest City in the Legend and Differ Across Plots)

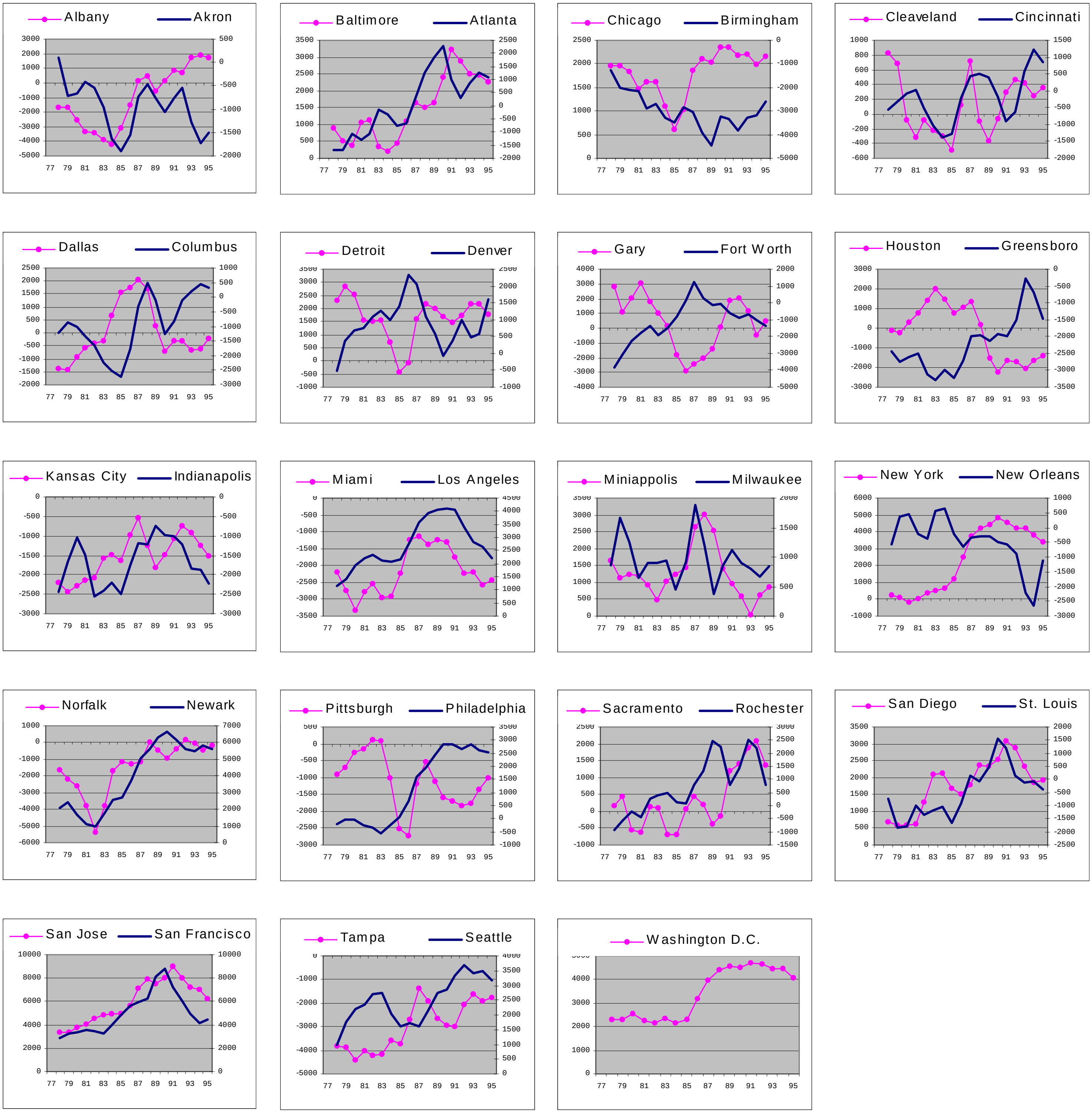

Figure 4: Worker Log-Population Shares - P^{worker}
(Vertical Scales Correspond to the Closest City in the Legend and Differ Across Plots)

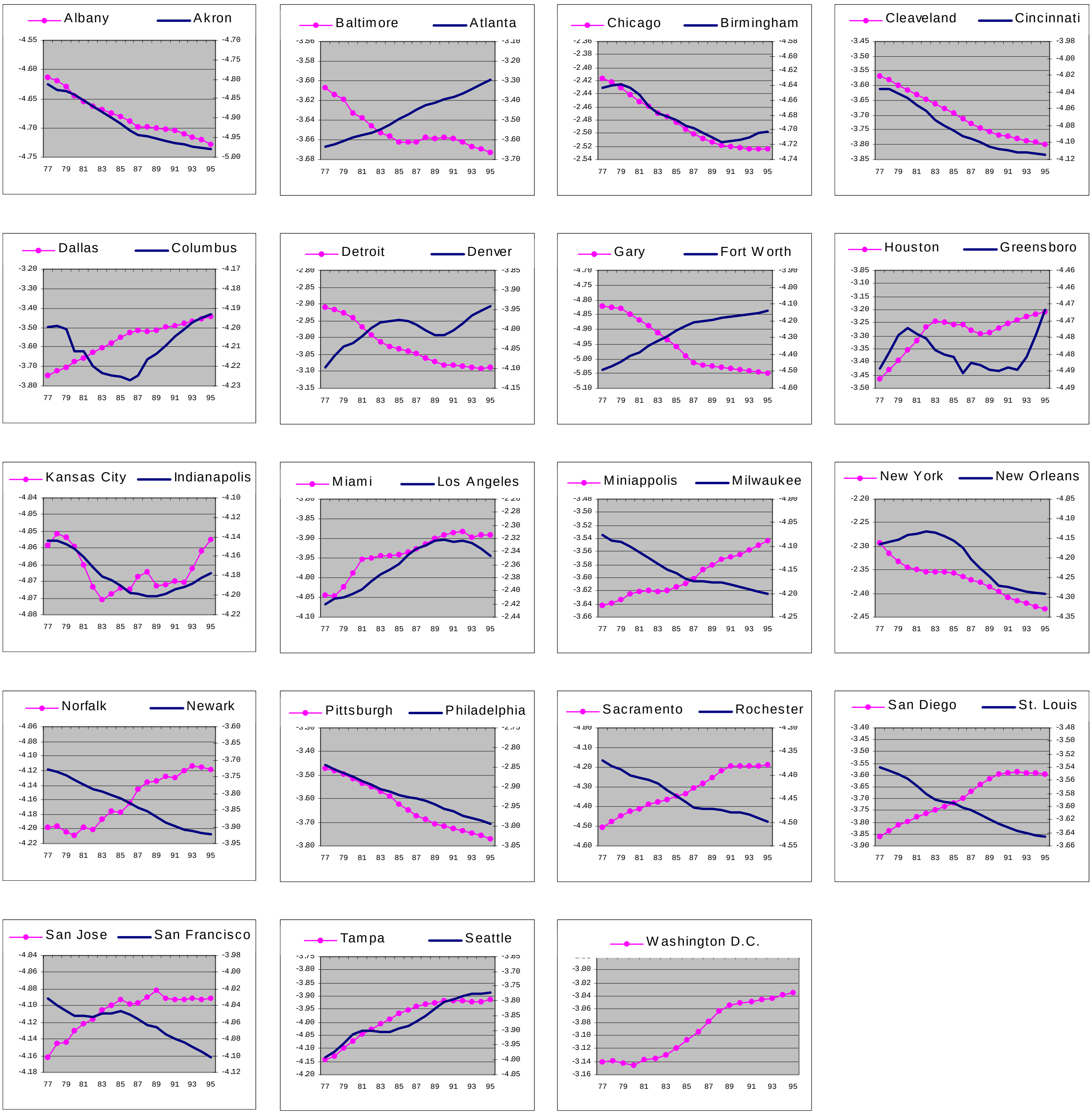

Figure 5: Retiree Log-Population Shares – $P^{retiree}$
(Vertical Scales Correspond to the Closest City in the Legend and Differ Across Plots)

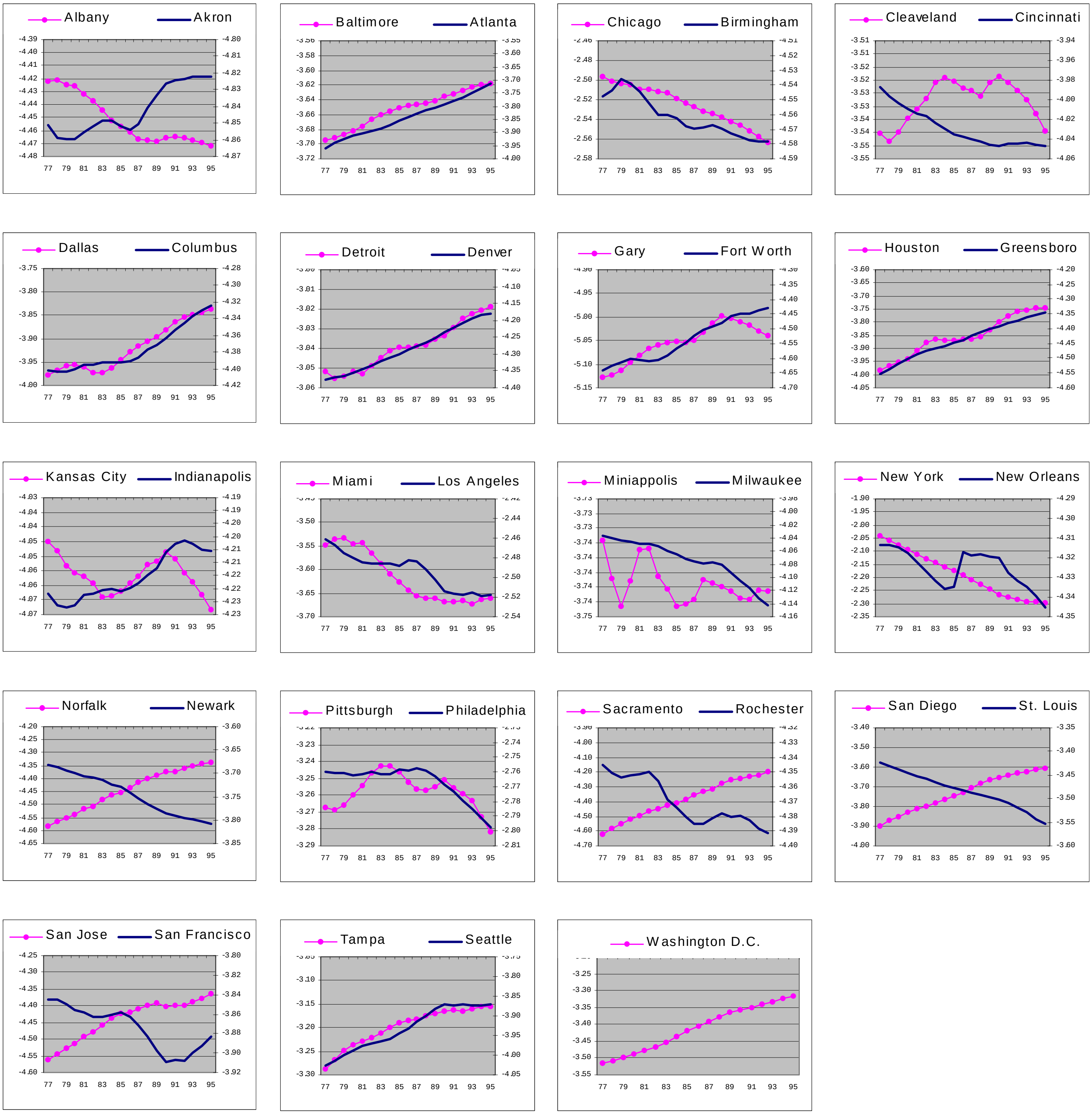

TABLE 2
ADF Unit Root Tests on Log-population shares, Q_H, Q_F, and r

	$\log(P^{worker})$	$\log(P^{retiree})$	$\log(P^{retiree}) - \log(P^{worker})$	Q_H	Q_F	r
Akron	-0.670	-2.751	0.604	-0.727	-2.042	0.141
Albany-Schenectady-Troy	-3.512 *	-3.717	-5.586	-2.612	-2.718	-2.212
Atlanta	-5.826 ***	-0.644	-3.576*	-2.762	-1.938	-1.504
Baltimore	-4.162 **	-2.294	-4.265**	-2.838	-1.840	-2.139
Birmingham	-0.311	-2.006	-0.749	-1.973	-0.608	-2.020
Chicago	0.029	-0.916	1.531	-1.213	-2.326	-1.988
Cincinnati	-1.228	-0.494	-3.614**	-3.059	-4.630***	-1.767
Cleveland-Lorain-Elyria	-1.063	-0.246	1.080	-1.908	-3.739**	-1.437
Columbus	-3.299	-1.090	-5.196***	-2.460	-3.365*	-1.354
Dallas	-1.723	-2.823	-4.188**	-1.664	-1.734	-1.900
Denver	-2.064	-1.609	-1.630	-3.150	-1.949	-2.159
Detroit	-0.614	-2.330	-0.193	-1.669	-2.158	-1.198
Fort Worth-Arlington	-1.399	-2.770	-3.388*	-0.879	-1.099	-1.897
Gary	-1.169	-0.848	0.582	-1.738	-1.240	-0.477
Greensboro--Winston-Salem--High Point	0.353	-2.332	1.361	-1.526	-2.001	-3.143
Houston	-3.286	-2.840	-3.341*	-1.595	-2.429	-2.967
Indianapolis	-0.876	-2.250	-0.019	-0.993	-1.608	-0.851
Kansas City	-1.343	-1.851	-0.242	-1.957	-1.368	-2.508
Los Angeles-Long Beach	0.689	-3.028	0.032	-3.596	-1.077	0.506
Miami	-2.947	-0.758	1.454	-1.803	-1.379	-2.167
Milwaukee-Waukesha	-2.716	-0.467	-0.679	-0.719	-4.024**	-1.036
Minneapolis-St. Paul	-1.770	-3.025	-2.553	-1.434	-1.787	-1.768
New Orleans	-3.798 **	-1.996	-4.054	-2.677	-3.404*	-2.039
New York	-3.284	-0.056	0.068	-1.894	-1.141	-1.339
Newark	-2.404	-3.690 *	-3.130	-2.376	-1.254	-1.852
Norfolk-Virginia Beach-Newport News	-1.532	1.083	-3.483 *	-3.315 *	-2.944	-1.603
Philadelphia	-3.145	-0.577	-0.072	-2.353	-2.284	-2.337
Pittsburgh	-1.712	-1.240	0.179	-3.751 **	-2.448	-1.656
Rochester	-1.831	-2.232	-3.114	-2.761	-2.431	-1.357
Sacramento	-3.085	-1.435	-3.539 *	-2.562	-2.212	-1.624
St. Louis	-1.870	-1.742	0.052	-1.599	-0.736	-2.048
San Diego	-2.456	-1.137	-2.774	-2.518	-2.366	-1.197
San Francisco	-2.321	-4.167 **	-1.086	-1.407	-1.202	-0.435
San Jose	-2.135	-2.082	-3.106	-2.748	-0.066	-1.096
Seattle-Bellevue-Everett	-3.378 *	-2.118	-2.268	-2.114	-1.892	-1.985
Tampa-St. Petersburg-Clearwater	-1.262	-1.151	-1.343	-2.020	-2.326	-1.983
Washington DC	-1.832	-1.641	-2.089	-2.446	-1.426	-2.990
Mean	-1.764	-1.685	-2.130	-2.130	-2.032	-1.659
Median	-1.851	-1.630	-2.020	-2.020	-1.949	-1.768
Critical Values for ADF Test	1% (***)	5% (**)	10% (*)			
from MacKinnon (1991)	-4.671	-3.735	-3.309			

TABLE 3
Error Correction Models With Common Long Run Effects Across Cities[a]
(t-ratios in parentheses)[b]

		Worker Log-Population Share	**Retiree Log-Population Share**		**Retiree Log-Population Share - Worker Log-Population Share[d]**	
First Stage Cointegrating Equation [c]	Q_H	1.18E-05 (8.215)	-2.61E-06 (-2.565)	1.56E-07 (0.226)	-1.45E-05 (-8.592)	2.65E-06 (2.319)
	Q_F	1.43E-05 (12.189)	-2.76E-06 (-3.332)	-	-1.71E-05 (-12.462)	-
	r	-	-	-5.52E-06 (3.332)	-	-3.42E-05 (-12.462)
Second Stage Error Correction Equation [c]	e_{t-1}	-0.23840 (-9.571)	-0.31609 (-11.206)	-0.31609 (-11.206)	-0.19104 (-8.730)	-0.19104 (-8.730)

[a]The estimated coefficients on Q_H, Q_F, and r in the cointegrating equations were constrained to be alike across cities. The coefficients on all of the other variables included in (5.1) and (5.2) were allowed to vary across cities. Those coefficients total 444 parameters over the two equations and 37 cities and are not presented to conserve space.

[b]A modified form of the Engle-Granger two-step method (Engle and Granger (1987)) was used to estimate the ECM to ensure that the standard errors and related t-ratios in the cointegrating equations have standard limiting distributions. Specifically, the cointegrating equations were estimated by ordinary least squares including 1 lead and 1 lag of the first differences of the two right-hand side variables of interest – either Q_H and Q_F or Q_H and r depending on the specification of the model (see Saikkonen (1991)).

[c]The dependent variables in the cointegrating equations are in log-levels. The dependent variables in the error correction equations are in log-first differences (annual growth rates of population shares).

[d]Retiree log-population share less worker log-population share equals $\log(R_j / N_j) - k$, the ratio of retirees to workers in city j less a constant k, where k equals the log-ratio of retirees to workers for the entire system of cities.

Appendix
Supplemental Tables

TABLE A-1[a]
Unit Root Tests on Log-Population Shares for Six Cities
Using Decade-by-Decade Data from 1800 to 1970

City	ADF Stat	City	ADF Stat
Baltimore	-0.724	Philadelphia	-1.532
Boston	-2.427	Pittsburgh	-.7552
New York	-2.063	Washington D.C.	-3.406*

[a]The symbols *, **, and *** indicate rejection of the null hypothesis of a unit root at the 10, 5, and 1 percent levels, respectively. All augmented Dickey Fuller (ADF) tests included two lags of the dependent variable. All tests included a constant and linear deterministic trend. Critical values for the ADF tests were from MacKinnon (1991) and equal -4.731, -3.761, and -3.323 at the 1, 5, and 10 percent levels, respectively.

TABLE A-2
Engle-Granger Cointegration Tests Based On
ADF Unit Root Tests on Residuals From the First Stage Cointegrating Equations

	$\log(P^{worker})$	$\log(P^{retiree})$	$\log(P^{retiree}) - \log(P^{worker})$
Akron	-2.375 **	-1.490	-2.389 **
Albany-Schenectady-Troy	-2.266 **	-1.824 *	-2.591 **
Atlanta	-3.589 ***	-2.966 ***	-2.919 ***
Baltimore	-2.113 **	-2.413 **	-1.943 *
Birmingham	-0.825	-2.630 **	-2.198 **
Chicago	-1.884 *	-1.682 *	-1.860 *
Cincinnati	-1.877 *	-1.424	-2.475 **
Cleveland-Lorain-Elyria	-1.312	-1.634 *	-1.003
Columbus	-2.393 **	-1.196	-2.508 **
Dallas	-1.399	-2.663 **	-2.008 **
Denver	-3.299 ***	-1.902 *	-3.954 ***
Detroit	-3.176 ***	-2.215 **	-3.800 ***
Fort Worth-Arlington	-1.277	-3.466 ***	-1.920 *
Gary	-1.031	-1.704 *	-1.024
Greensboro—Winston-Salem--High Point	-1.666 *	-1.445	-1.427
Houston	-2.656 **	-2.901 ***	-2.368 **
Indianapolis	-0.946	-1.757 *	-1.235
Kansas City	-3.027 ***	-2.231 **	-1.880 *
Los Angeles-Long Beach	-4.654 ***	-2.038 **	-3.197 ***
Miami	-2.805 ***	-0.911	-0.666
Milwaukee-Waukesha	-2.359 **	-1.858 *	-1.773 *
Minneapolis-St. Paul	-2.236 **	-3.436 ***	-3.252 ***
New Orleans	-2.281 **	-2.276 **	-2.244 **
New York	-1.747 *	-2.252 **	-3.366 ***
Newark	-1.263	-1.924 *	-1.207
Norfolk-Virginia Beach-Newport News	-3.018 ***	-1.030	-0.617
Philadelphia	-2.088 **	-1.062	-1.474
Pittsburgh	-1.115	-3.091 ***	-1.263
Rochester	-1.513	-3.008 ***	-1.439
Sacramento	-2.560 **	-3.523 ***	-7.744 ***
St. Louis	-1.957 *	-1.535	-1.585
San Diego	-2.142 **	-2.114 **	-2.791 ***
San Francisco	-1.946 *	-2.068 **	-2.224 **
San Jose	-2.651 **	-2.315 **	-2.850 ***
Seattle-Bellevue-Everett	-2.341 **	-2.584 **	-1.797 *
Tampa-St. Petersburg-Clearwater	-1.410	-1.059	-1.617
Washington DC	-3.534 ***	-1.705 *	-2.127 **
Mean	-2.182 **	-2.090 **	-2.236 **
Median	-2.142 **	-2.038 **	-2.008 **
Critical Values for ADF Test	1% (***)	5% (**)	10% (*)
from MacKinnon (1991)	-2.757	-1.968	-1.629