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Mark D. Partridge
AED Economics, The Ohio State University

Dan S. Rickman
Oklahoma State University

M. Rose Olfert
University of Saskatchewan

Ying Tan
Oklahoma State University

Department of Economics
Oklahoma State
University
Stillwater, Oklahoma

339 BUS, Stillwater, OK 74078, Ph 405-744-5110, Fax 405-744-5180

When Spatial Equilibrium Fails:
Is Place-Based Policy Second Best?

by

Mark D. Partridge [1]
Dan S. Rickman[2]
M. Rose Olfert [3]
Ying Tan[2]

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Abstract: Place-based or geographically-targeted policy has been promoted as a way to help poor regions and the poor people who live there. Yet, such policy has often been attacked by economists as slowing needed economic adjustments, redirecting resources to lower productivity regions, and supporting political agendas rather than economic prosperity. The spatial equilibrium model in particular predicts that people readily move to the locations providing the highest expected utility, suggesting that policy interventions only impede needed adjustments. Given the high mobility of Americans, the spatial equilibrium model should then be most applicable to the US. We review the empirical evidence on whether the spatial equilibrium model applies and find that, even in the United States, people are not as mobile as the model suggests and that economic shocks have rather persistent effects. Although this suggests the potential need for place-based policy, we describe the informational and political economy conditions that need to be met before place-based policy can be effective.

1. AED Economics, The Ohio State University, email: partridge.27@osu.edu
2. Department of Economics and Legal Studies, Oklahoma State University, Rickman email: dan.rickman@osu.edu; Tan email: yingt@ostatemail.okstate.edu.
3. School of Public Policy, University of Saskatchewan, rose.olfert@usask.ca

1. Introduction

The conventional spatial equilibrium view of regional economies—at least in its strong form—precludes the need for place-based policies. According to this view, utility differentials, such as those created by regionally-asymmetric demand shocks, induce internal migration. Migration arbitrages away the utility differentials such that relative employment and real wage rates return to their equilibrium levels, capitalizing the attributes underlying the utility differentials into factor prices (Partridge and Rickman, 2003a). The question is the degree to which this view accurately describes the workings of regional economies, and if so, do adjustments occur in a timely manner, obviating the need for place-based policy.

In this paper we provide a thorough review of the empirical literature regarding the efficacy of the equilibrium view of regional economies.[1] We then update the empirical literature in testing the mechanisms and success of spatial equilibrating forces. Finally we provide a Discussion of the policy implications, specifically addressing the conditions under which place-based policy may be warranted.

2. Literature Review

Because it is thought to possess the most flexible labor market, with historically high relative migration rates, most of the literature and our review pertain to the US economy. Related discussion for other countries is presented when appropriate.

The literature generally consists of three broad strands. First, a number of cross-sectional tests have been conducted regarding equalization of household utility and price-adjusted wages across space. Because of the difficulty in assessing the utility value of regional attributes at a point in time, a potentially more fruitful way to test spatial equilibrium is the examination of regional labor market responses to shocks in the medium and long run (Glaeser and Gottlieb, 2008). Thus, a second strand examines the persistence of differentials in key regional labor

[1]Our review will be at the larger regional level—say at the US county-level. In our extensive review, we will not review place-based (also called geographically targeted) policies at the neighborhood or other small areas such as enterprise zones. For a review see Hansen and Rohlin (2011).

market indicators. Yet, this literature provides only indirect evidence on the research question regarding the potential need for place-based demand policies because adjustments to clearly identifiable demand shocks are not examined. A third strand examines directly the effects of regional employment growth on labor market outcomes. Because employment growth can be driven by either demand or supply forces, many studies use the shift-share model to derive an exogenous demand-based measure of employment growth.

2.1 Cross-Sectional Tests of Spatial Equilibrium

A fundamental result of the spatial equilibrium model is that differences in incomes do not necessarily reflect utility differentials and, alone, cannot be used to justify place-based policies. Income differences may reflect many other factors such as differential amenity attractiveness of areas. Costs of living such as housing costs also may differ, offsetting nominal income differences. To be sure, income and housing prices capitalize the values of locational attributes in spatial equilibrium. Property owners, rather than working households in poor areas, may be the primary beneficiaries of places receiving financial assistance (Glaeser and Gottlieb, 2008). Therefore, tests of spatial equilibrium have examined differences in price-adjusted wages and subjective measures of utility.

An extensive early literature exists on whether price-adjusted wages are equalized across regions, generally finding that they are not (Dickie and Gerking, 1989). However, the spatial equilibrium model posits that utility, not price-adjusted-wages, will be equalized. Lower price-adjusted wages may reflect the existence of household amenities. Thus, subsequent studies attempt to ascertain whether household utility is equalized across regions.

Because utility is not directly observable, tests of utility equalization examine price-adjusted wages while attempting to control for differences in natural amenities. Assuming marginal differences or a Cobb-Douglas utility function, the elasticity between wages and prices across regions should be unity for utility to be equalized (Winters, 2009). Roback (1988) reports an elasticity of 0.97; using a different price index, DuMond et al. (1999) find a wage-price elasticity of 0.46. Winters (2009) attributes the difference in results at least partly to the use of different price deflators in the two studies. Winters finds this elasticity to depend upon the housing price measure and method of estimation. Using housing values and OLS, he finds the elasticity to be less than 0.5, which rises to 0.76 when housing rents are used. Including housing rents, and employing instrumental variable estimation to account for measurement error, leads to an estimated elasticity not statistically significantly different from unity, supporting the existence of a spatial equilibrium.

The estimated wage-price elasticity then likely depends greatly on the price deflator, estimation method, and amenity measures included in the regression. Thus, more direct, self-reported measures of utility have been examined, in which the spatial equilibrium approach suggests that utility and per-capita income are not necessarily related. For example, using the General Social Survey between 1972 and 2006, Glaeser and Gottlieb (2008) find that the fraction of respondents in US metropolitan areas who report being happy was unrelated to income per capita, though there were large differences in reported happiness across areas.

Oswald and Wu (2011) use self-reported responses from surveys conducted by the US Center for Disease Control as part of the Behavioral Risk Factor Surveillance System to examine life-satisfaction across states. They do not find any significant correlation between self-reported life-satisfaction and GDP per capita, though after controlling for income, the correlation becomes strongly negative, which the authors interpret as evidence in favor of the weak form of the compensating differential hypothesis. Yet, life-satisfaction differences remained after controlling for individuals' backgrounds and characteristics, leading the authors to reject the strong form of the compensating differential hypothesis that utility levels are equalized across states.

Rickman (2011) reports Oswald and Wu's state ranking of residents' average life-satisfaction level to be significantly correlated with a state population growth ranking for the period 2000 to 2010 (r=0.48), suggesting disequilibrium adjustment. Clark et al. (2003) derive measures of over- and under-compensation in local labor markets and use them to examine interregional migration. They find net migration towards areas where there is estimated over-compensation and away from areas where there is under-compensation. Migration, thus, worked to arbitrage away compensation differentials, consistent (at least weakly) with spatial equilibrium theory. Yet, over- and under-compensation in regional labor markets represent disequilibria during the study period.

Bayer et al. (2009) find that labor mobility costs preclude environmental amenities from becoming fully capitalized into factors prices. Labor mobility may be hindered because of household ties to the area and home ownership, leading hedonic analysis to undervalue environmental amenities.[2] Greenwood et al. (1991) also find that the assumption of spatial equilibrium causes understatement of the equilibrium values of compensating differentials. Using the reported estimated disequilibrium gaps for 1980 by Greenwood et al. (1991), we find that when controlling for state natural amenity attractiveness, there are significantly positive relationships with population growth over both the subsequent five and ten years.[3,4]

2.2 Persistence of Regional Labor Market Differentials

Numerous studies have examined the reasons for regional labor market differentials and their persistence over time. Many have focused on disequilibria in regional labor markets arising from differences in labor demand, while others emphasize equilibrium explanations such as amenities and labor market policies (Holzer, 1993; Partridge and Rickman, 1997a). Because of the attention given to unemployment at the national level, regional differentials in unemployment have received the most attention, with some attention given to regional differentials in employment rates and poverty. For place-based policymaking, the most important aspect of regional labor market differentials is the persistence of the disequilibrium component. Early empirical investigations relied on econometrically-estimated partial adjustment models, while later investigations used sophisticated time-series methodology, which included controlling for

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[2] Partridge and Rickman (1997a) find that a greater percentage of the population born in the state and greater home ownership (as measures of mobility costs) led to persistently higher US state unemployment rates.

[3] Estimates of state migration are derived from the US Census Bureau as the residuals of the change in population less natural population increases (http://www.census.gov/popest/data/historical/1980s/state.html, accessed May 14, 2012).

[4] The respective R^2s are 0.85 and 0.75, in which the R^2s for the disequilibrium differentials used alone decreased from 0.57 to 0.48. This is suggestive of persistence in disequilibrium that diminished in the longer run, though in the longer run there is a greater chance of additional demand shocks also influencing migration.

the possibility of shifts in equilibrium differentials.

Marston (1985) used a variance components method to assess the relative sizes of equilibrium and disequilibrium components of unemployment for a sample of US metropolitan areas during the 1970s. He reported that the disequilibrium component was only half as large as the equilibrium component and that it did not last for more than a year. Using a similar approach with post-WWII data through 1992, Davis et al. (1997) finds much more persistence in unemployment, with the magnitudes of the coefficients on lagged unemployment suggesting that only 30 percent of unemployment differential is eliminated in one year. The difference in findings is argued to relate to Marston's (1985) use of a short panel period. They report negative shocks as having a larger absolute value effect on unemployment than positive shocks.

Based on an examination of US metropolitan areas from 1976 to 1984, Hyclak (1996) finds that about a one point rise in unemployment is reduced to one half by the end of four years. Gordon (1988) finds London's unemployment differential to dissipate fairly quickly following the initial rise in response to a drop in employment. Yet, the dissipation of the residual was slower during times of high national unemployment.

Partridge and Rickman (1997b) decomposes US state unemployment differentials into equilibrium and disequilibrium components for 1992-1994. They report that for some regions, the disequilibrium component was larger, while in other regions the equilibrium component dominated. They do not address the issue of persistence in the disequilibrium component, though contemporaneous and lagged regional employment growth measures are found to be significantly related to regional unemployment.

Migration flows have also been examined in the context of interregional migration being sufficient to eliminate disequilibrium unemployment differentials. Marston (1985) concludes that the flows of people between areas in a year, is large compared to the disequilibrium component. In examining US place-to-place migration flows during the 1980s, Gabriel et al. (1993) conclude they were insufficient to offset shocks to the regional distribution of unemployment rates and that the primary effect was on wage differentials. Treyz et al. (1993) estimate net migration for

US states, reporting a sluggish migration adjustment process to regional imbalances in demand. They find that it took twenty years for 83.5 percent of a demand-induced imbalance to be eliminated through migration. Using US state data from 1976-1996, modeling migration as forward looking, Gallin (2004) concludes that all of the migration adjustment occurs within ten years.

Pissarides and McMaster (1990) estimate migration and wage pooled cross-section regressions for British regions and link them to an equation for unemployment. Based on simulation of the three equations, they conclude it would take more than twenty years to eliminate a disequilibrium unemployment differential. Groenwald (1997) uses three econometric equations for wages, unemployment and migration for an Australian State, reporting that adjustment through interregional migration is slow—12.5 percent of a disequilibrium unemployment differential remained after fifteen years.

More recently, studies of unemployment persistence have shifted from econometric equations to more formal time series tests. In particular, unit root tests have been performed to determine whether shocks to regional unemployment have persistent effects. A unit root is said to exist if one fails to find evidence that the coefficient on the lagged unemployment rate is less than one. A unit root indicates persistence because a value of one indicates that regional unemployment will not revert back to the mean or trend value following a demand shock; i.e., shocks have permanent effects. Values less than, but near, one still imply sluggish adjustment.

Blanchard and Katz (B&K) (1992) perform Augmented Dickey-Fuller (ADF) unit root tests for US state unemployment over the period 1972-1990, failing to reject a unit root in unemployment in all but two states. Because of the known low power of the test to reject the null hypothesis of a unit root in small samples, and based on theoretical prior beliefs of regional unemployment rates, B&K model them as stationary. Stationarity of unemployment implicitly forces migration to arbitrage away regional unemployment differentials (Obstfeld and Peri, 1998). In comments on the B&K paper, Robert Hall (1992) notes the short time series on which their findings are based. Rowthorn and Glyn (2006) use a longer employment time series that contains less measurement error and find that employment rates permanently change in response to shocks, indicating that migration plays a much smaller role in state labor market equilibration than concluded by B&K.

Recognizing the low power of the unit root tests in small samples, Payne et al. (1999) confirm the ADF results with the use of variance ratio tests. They fail to reject the null of a unit root in unemployment for all fifty US states during 1978-1996, supporting the view that shocks to state and aggregate unemployment have permanent effects. In addition, for only two states do they find co-integration between the state and national unemployment rates, indicating that the regional differentials are nonstationary. This suggests persistent disequilibrium differences in regional unemployment rates. In contrast, in an analysis of UK regional unemployment rates for 1965-1995, Martin (1997) finds them to be co-integrated with the national rate. He estimates that divergences from long-run equilibrium are eliminated within four to six years.

Subsequently, newer generations of unit root tests have been employed to examine regional unemployment rates. Song and Wu (1997) test for unit roots in unemployment rates for the forty-eight contiguous US states. While they fail to reject the unit root for most states using ADF and Phillip-Perron tests, the null of a unit root is decisively rejected using a more powerful panel-based (Levin-Lin-Chu) test that imposes cross-section restrictions, casting doubt on the existence of hysteresis in state unemployment rates. Yet, a null hypothesis of the coefficient on the lagged dependent equaling zero also is rejected, indicating persistence in state unemployment rates.

Unit roots are less likely to be rejected when the time series possess structural breaks (Bayer and Jüßen, 2007; Romero-Ávila and Usabiag, 2007; Sephton, 2009). Nevertheless, the question of persistence remains. Romero-Ávila and Usabiag (2007) find the average half-life of impulse responses to shocks in US state unemployment to be six years, with a convergence speed of 11 percent per year, in which the upper bound of the confidence interval exceeded twenty years in all but nine states. Sephton (2009) suggests that unemployment persistence diminishes after the second break that typically occurred in states around 2001. Bayer and Jüßen (2007) in examining West German regional unemployment rates find that allowing for a structural break reduces the half-life of a shock from 5.6 to less than two years on average, concluding that small government interventions will not likely be effective, though a regime shift suggests that a large intervention could move the economy from one equilibrium to another.[5]

In continuing developments in the literature, panel unit root tests that impose cross-sectional independence, where there is dependence, produce over-sized tests. Therefore, León-Ledesma (2002) demonstrates the importance of addressing cross-sectional dependence in panel unit root testing of unemployment rates for US states and EU countries, failing to reject the unit root for most states and countries, and finding slower adjustment in EU countries. Cheng et al. (2012) provide the most recent evidence on US state unemployment persistence. They find evidence of nonstationarity when the most recent recession is included and even where there is evidence of stationarity they find the half-life of a common component ranging from 6 to 14 years.[6]

Although unemployment has garnered most of the attention in the literature, some attention also has been given to the persistence of regional poverty. In a series of studies using a partial adjustment model, Partridge and Rickman find US regional poverty to be very persistent (Partridge and Rickman, 2005; 2006b; 2007a; 2007b; 2008a; 2008b), in which the fraction of the regional poverty differential that is eliminated over a ten-year period ranges from 40 to 70 percent. The poverty adjustment process in high poverty nonmetropolitan counties, including persistently high poverty counties, is no more sluggish than that of other nonmetropolitan counties (Partridge and Rickman, 2005; 2007a). Metropolitan areas are found to have poverty at least as persistent as that for nonmetropolitan areas (Partridge and Rickman, 2008b). Using several unit root tests, including panel unit root tests, Wang and Dayanandan (2006) find that generally the tests support nonstationarity of poverty for most Canadian provinces over the period 1980 to 2003.

[5]Gomes and da Silva (2009) use an endogenous one/two-break LM unit root test and find that the unit root null hypothesis cannot be rejected for the major metropolitan areas of Brazil, with the exception of Rio de Janeiro. They conclude that the high persistence in Brazilian regional unemployment rates that would be difficult to overcome.
[6]Yet, Garcia-del-Barrio and Gil-Alana (2009) find that regional unemployment is persistent in Spain using panel unit root tests regardless of whether they account for cross-sectional dependence.

2.3 Employment Growth Effects on Regional Labor Market Outcomes

An extensive number of studies directly examine the effects of differentials in employment growth on regional labor market outcomes. While some simply use ordinary least squares of single econometric equations or reduced-form VAR equations, a common approach is to use the industry mix component from regional shift-share analysis in instrumental variable estimation. The industry mix component is simply the growth in employment that would occur in a region if all of its industries grew at their corresponding national rates. In addition to capturing the direct effects of having fast- or slow-growing industries, because of multiplier effects, the measure also captures some growth in industries that differ from their national averages. So long as industries are not concentrated in single regions and a region's composition of industries is exogenous to regional labor market outcomes, the measure provides a useful exogenous instrument to assess regional labor market effects of labor demand shocks. In addition, the studies range from examining employment-induced responses in unemployment and labor force participation to the employment growth effects on the distribution of regional income.

2.3.1 Vector Autoregression Studies

Based on estimated reduced-form VARs for individual states, B&K find that following a shock to employment, state unemployment and labor force participation rates return to their previous levels on average after five to seven years. However, pooling all states for 1978-1990, using longer lags, and various measures (instruments) of (for) employment growth, they find that about 15-17 percent of the unemployment and labor force participation responses to an employment shock remain after eight years. Bartik (1993a) argues that large measurement error inherent in the unemployment rate and labor force participation and restrictions on lag lengths bias B&K's results towards no long-run effects. Using the same data as B&K, but testing for the optimal lag length rather than using B&K's restriction of two years used in the single equation VARs, Bartik finds that after seventeen years, 25% of an employment shock is still reflected in the regional labor force participation rate.

Decressin and Fatas (1995) report similar findings for the US, but find that in contrast to

Europe, labor force participation absorbs most of the labor demand shock. In contrast to the findings for the US and Europe at the time, Jimeneo and Bentolila (1998) find that about one-fourth of a labor demand shock effect on regional unemployment and labor force participation in Spain remains in the long run. They also report that the Spanish responses are slower than for the US and the rest of Europe. Using the B&K approach, Fredriksson (1999) reports that regional adjustment to labor demand shocks in Sweden is rapid, in which interregional migration is the primary adjustment and unemployment and labor force participation rates return to normal within two years. Also using the B&K approach and assumptions to examine 166 regions in Europe for 1988-1997, Tani (2003) concludes that European workers are much more mobile than previous studies suggested, though as already noted, there are reasons to believe that B&K's methodology understates persistence and overstates mobility.

Hunt (2006) finds US employment and population to be co-integrated. Based on estimated vector error correction models, 53 percent of the impulse responses took place within 20 years, with 73 percent taking place within 25 years. These, he argued, were significantly shorter than those estimated in previous studies that used levels of nonstationary variables. Yeo, Sung and Holland (2005) examine employment, population, and labor force participation for the state of Washington. All three variables are found to contain a unit root and are co-integrated. A shock to employment has a permanent long-run effect on labor force participation; almost 30 percent of the initial effect on labor force participation remains in the long run.

Partridge and Rickman (2003b; 2006a; 2009) construct long-run restrictions structural vector autoregression (SVAR) models to assess regional labor market dynamics. An advantage of the long-run restrictions approach is that employment is no longer assumed to solely represent labor demand as in B&K and other reduced-form VAR studies. Rather, explicit recognition is given to the roles of both labor demand and labor supply; migration contains labor supply innovations and does not simply represent a response to labor demand. Among the primary findings, about 20 percent of a labor demand shock is reflected in the employment rate in the long run for the United States (Partridge and Rickman, 2006a), varying from a low 13 percent for

Sunbelt states to 55 percent for Rustbelt states. For Canada, about one-third of a labor demand shock is reflected in the employment rate in the long run, with larger estimates for Ontario and Quebec, with little effect found for the Atlantic Provinces (Partridge and Rickman, 2009).

2.3.2 Econometric Equation Studies of Unemployment and Labor Force Participation

One approach to examining labor supply responses to employment growth is to decompose the change in employment into the sources of supply using an identity. The supply responses then are regressed on employment growth; because of the identity, the sum of the estimated supply responses equal unity. Two primary assumptions underlie the approach: 1) short-run fluctuations in employment are demand-driven; 2) labor supply contemporaneously responds to labor demand shocks without a lag (Partridge and Rickman, 2008d).

Eberts and Stone (1992) use this approach in examining changes in unemployment, labor force participation and population for US metropolitan areas. They find increased labor force participation of local residents to be the primary labor supply response to increased job growth (pp. 23-24), where three-fourths of a change in employment is satisfied by individuals entering/ exiting the labor force. Changes in the unemployment rate and population share account equally for the remaining supply adjustments. Across Census regions, the population share ranges from a high of 22.8 percent in the Mountain region to a low of 2.1 in the West North Central region. In subsequent analysis, using recursive identifying restrictions in estimating regional labor demand and supply, Eberts and Stone (1992, pp. 78-83) conclude that the time for full adjustment of local labor markets toward a new equilibrium following a demand shock exceeds a decade.

Because they examined US counties, a finer level of disaggregation, Partridge and Rickman (2008d) added net commuting as a fourth potential source of labor supply. Based on their decomposition, they conclude that generally net commuting was the dominant supply response, limiting the benefits of job growth to the original residents. However, in nonmetropolitan counties, a change in labor force participation was the primary response. Persistently high-poverty nonmetropolitan counties had the smallest commuting response and largest labor force participation and unemployment responses.

In comparing London to other areas in Great Britain, Gordon (1988) found that the unemployment rate response to a change in employment varied by size and type of area and national economic conditions. Only 17 percent of employment-induced change to unemployment remained after one year, where the largest response was migration. In Scotland (taken as a region), over 40 percent of the unemployment effect remained after a year, whereas, in the typical inner borough of London, all of the effect was eliminated in one year.

In a review of the early literature, Bartik (1991, Ch. 4) reports that most studies only estimated employment growth effects for short periods of time, and where longer periods were considered (10 years or greater), the studies did not distinguish short-run from long-run effects. He finds that one-percentage point employment growth effects on regional unemployment rates over long periods of time ranges across studies from -0.04 percentage points to no effect, while for labor force participation they range from no effect to 0.08 percentage points. The studies also failed to distinguish between whether employment growth represented demand or supply, whether the effects varied across areas by growth rates or across types of individuals.

In an examination of US metropolitan areas, Bartik (1991) finds that a one-percentage point increase in employment growth reduces unemployment by 0.06-0.07 percentage points in the long run, depending on whether micro or aggregate data are used. The labor force participation rate is increased by 0.14 percent. He finds generally similar effects across groups of people, though labor force participation rates of older workers are significantly more sensitive to local employment growth.

A number of studies also examined whether employment growth effects on unemployment and labor force vary by region or by source of employment growth. Bartik (1991, Ch. 4) finds no difference across metropolitan areas experiencing different growth rates. Using pooled data for 38 US metropolitan areas, Bartik (2009) finds that regional employment growth’s effect on employment rates is significantly higher in the short run for metropolitan areas with initially slacker labor markets but generally not in the longer run. Yet, there is some evidence that in areas with the tightest labor markets, employment effects are smaller than the average.

Partridge and Rickman (1997a; 1998) find that faster US state employment growth associated with a state's composition of nationally fast-growth industries reduces unemployment rates more than growth idiosyncratic to the state. They attribute this to lesser migration in response to industry mix employment growth because if the industries were growing faster nationally, and there was imperfect labor mobility across sectors, there would less incentive for migration. This result also was reported in the migration study of Partridge and Rickman (1999).[7] Partridge and Rickman (1998) found that only idiosyncratic employment growth reduced the short-term unemployed share of the labor force more than the long-term share because of competition of migrants with the long-term unemployed. This is in contrast to industry mix employment growth with a general lack of migration response.

Using an industry mix employment growth measure of labor demand shocks for US metropolitan areas, Notowidigdo (2011) finds that positive shocks increase population and employment more than negative shocks reduce them. This asymmetry is particularly true for low-skilled workers. The reasons attributed to the lower mobility in declining regions are a lower elasticity of housing supply because housing stock is durable (Glaeser and Gyourko, 2005) and social transfer payments. Because of the larger share of their budget spent on housing, low-skilled workers particularly benefit from falling housing prices in declining regions. Although social transfer payments limit labor mobility, Notowidigdo finds that even if the transfer payments were replaced by subsidies, the housing sector alone causes asymmetric population and employment responses to demand shocks in growing versus declining regions.

2.3.3 Econometric Equation Studies of Income Distribution Effects

Evidence of employment growth effects on regional labor market outcomes can also be found in regional studies of income and its distribution, including poverty. In a review and updating of previous research, Bartik (2005) concludes that five years after a one percent

[7]Partridge and Rickman (1999) find that a one percent increase in competitiveness employment increased migration by 0.29 percentage points, but industry mix employment increased it by only 0.14 percentage points. These are within the range reported by Bartik (1991), though on the low side, being slightly lower than the estimate of Treyz et al. (1993), which translate into a lengthy adjustment period within a stock adjustment model.

increase in local employment, there is an increase in real earnings per capita that is 0.28 percent of local area personal income. Half of this results from area residents moving up to higher-paying occupations (the other half was attributed to increased employment rates). In further reviewing the literature, Bartik (forthcoming) concludes that a one percent demand shock to local employment increases local employment rates by 0.2 percent and occupation wages by 0.2 percent, for a total effect on real earnings per capita of 0.4 percent. If the local job mix shifts towards industries that pay more nationally, there are additional wage spillover benefits. A one percentage point increase in wages because of a shift in the mix of high-paying jobs increases local earnings by two percentage points (Bartik 1993b).

Regarding the distribution of income gains from employment growth, in a study of US metropolitan areas from 1979-1988, Bartik (1994) finds that employment growth in a metropolitan area increases the share of income received by those in the lowest income quintile in the long run. Strong employment growth particularly benefits workers with the least skills and education because a tight labor market forces employers to hire them. Bartik cautions that they only benefit if economic development program costs are modest and not financed by cuts to social programs that benefit the poor.

Based on panel data for individuals in US metropolitan areas, Bartik (1996) finds that a boost to employment growth of one percentage point reduces the probability of poverty for females by 0.33 percent and males by 0.20 percent. He also finds the same increase in MSA employment growth as increasing average real earnings by 0.5 or 0.6%, in which half is attributable to increased annual hours worked, and the other half to greater wages. In contrast to other studies (e.g., Bartik, 1993b), blacks and whites experienced similar real earnings gains from increased MSA growth. Bartik (2001, p, 148) concludes that 10 to 20 percent of the increase in employment and earnings may persist in the long run, in which the most important channel is poor individuals moving into higher paying jobs.

Bound and Holzer (2000) examine the effects of labor demand shifts for US metropolitan areas during the 1980s. They find population responses partially offset the labor market effects of demand shifts. But the more limited population responses among less-educated workers lead to them experiencing greater losses in work hours and earnings. Based on instrumental variable estimates (using industry mix employment growth as the instrument), a 10 percent decline in an area's labor demand would reduce nominal earnings by 11 percent for high school educated workers and 6 percent for workers with a college degree. Most of these nominal wage declines are estimated to translate into real wage declines.

Using an equilibrium framework, Levernier et al. (2000) generally do not find recent employment growth to have reduced US county poverty in 1989. Yet, they find employment to reduce poverty in counties with larger shares of African-American residents. However, using a disequilibrium framework, Partridge and Rickman (2006a, p. 142) find a one percentage point increase in employment growth reduces US county poverty by 0.37 percentage points in 1989 and 0.23 percentage points in 1999.

In a study of US metropolitan areas (using a disequilibrium framework), Partridge and Rickman (2008b) find employment growth to reduce poverty, in which the effect varies across metropolitan size and county type. A one percentage point increase in employment growth reduces poverty by 0.2 percentage points in large metropolitan area (MA) central city counties in the short run and about 0.4 percentage points in the long run. The null hypothesis of exogeneity of poverty to employment growth using the instrument of industry mix job growth from shift-share analysis could not be rejected. No effect is found for large MA suburban counties. For medium-sized (small-sized) MAs, a one percentage point metropolitan-wide increase in employment reduces poverty by 0.3 (0.3) percentage points in the short run and 0.5 (0.6) percentage points in the long run. No county-level employment effects on poverty are found for multi-county medium-sized metropolitan areas.

Examining poverty rate changes in the 1990s for US Census tracts, Crandall and Weber (2004) show that job growth reduces poverty more in high-poverty neighborhoods. A one percentage-point increase in employment growth rates reduces poverty by 0.011, 0.046 and 0.088 percentage points in low, medium and high poverty tracts, respectively. The results are notable given the smaller geographic scale of Census tracts. They also find that a higher initial poverty rate in a tract was associated with a greater poverty decline over the subsequent decade.

Partridge and Rickman (2006b, Ch. 4) find an increase in current and lagged (up to two years) state employment growth of one percent reduces the poverty rate by 0.5 percent. They also present evidence that the effect is stronger during times of low national unemployment. The channels through which employment growth is found to affect poverty include reducing unemployment, increasing employment rates, and reducing teen birth rates.

For nonmetropolitan high-poverty US counties, Partridge and Rickman (2005) estimate that a one percentage point increase in employment reduces poverty by 0.11 percentage points in the long run, approximately double the magnitude of the estimate for other nonmetropolitan counties. When relative employment growth is decomposed into two components using shift-share analysis, the industry mix and competitiveness components have approximately the same poverty effect in high-poverty counties. Yet, for other counties, the poverty-reducing effect is four times larger for growth attributable to the county's mix of industries. Partridge and Rickman (2005) argue that the stronger industry mix effect in non-high-poverty counties occurs because of a lesser migration response if the industries are faring poorly nationally, consistent with evidence presented above for unemployment.[8]

In a follow-up study on persistent poverty counties, Partridge and Rickman (2007a) further examine the poverty generating process in persistently high-poverty nonmetropolitan counties using Geographically Weighted Regression. They find that employment growth has three-times the magnitude effect on poverty in persistently high-poverty counties relative to other nonmetropolitan counties. However, the effect is not found to vary across differing persistent-poverty county clusters.

In a pair of related papers, using Geographic Information Systems data, Partridge and Rickman (2008a; 2008c) find that remoteness influences how job growth affects poverty. Using

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[8] In a related study of persistently high-poverty US counties, Partridge and Rickman (2005) find that recent employment growth significantly increases the probability that a county moves out of high-poverty status.

industry mix employment as an instrument, local job growth reduces poverty only in nonmetropolitan counties at a sufficient distance from the nearest metropolitan area, presumably because of lower commuting and migration responses. In fact, both studies reveal lower migration responses to employment growth the farther a nonmetropolitan county is from a MA. They conclude that higher poverty in remote areas is not simply the result of the sorting of poor people into these areas, but a result of adverse labor demand shocks that do not dissipate through labor market adjustment (i.e., sorting implies that job growth would not help the self-sorted households that do not desire employment). In addition, job growth in the nearest MA is found to reduce poverty but the effect attenuates with distance.

3. Updated Empirical Test of Spatial Equilibrium

We now briefly report some updated empirical tests of spatial equilibrium for the US to illustrate the potential failures of the spatial equilibrium process to even out economic outcomes, raising the possibility that place-specific policy may be warranted. Given the high mobility rates of American households, US results could be viewed as near the upper bound of an efficacious spontaneous spatial equilibrating process. Our sample includes over 3,000 continental US counties including the District of Columbia.[9] Due to different expected rural and urban responses, we separate metropolitan area (MA) and nonmetropolitan area counties in the empirical analysis.[10] We focus on the 1990-2000 and 2000-2010 time periods for comparison; the former period a boom period and the latter characterized by general economic sluggishness with little job creation. The latter period also exhibits a general decline in economic migration (Partridge et al., 2012), which would reduce the natural economic equilibrating effects in local labor markets. We also consider the 2000-2007 and 2007-2010 periods to assess whether the Great Recession (which began in Dec. 2007) altered the longer-term patterns.

The dependent variables consist of measures related to economic outcomes associated with

[9]Following the US Bureau of Economic Analysis, there are cases where independent cities are merged with the surrounding county to form a more functional region (mostly in Virginia). Forty three mostly small rural counties are omitted due to the lack of economic data.

[10]Generally, we use the 2003 MA definitions. See the US Census Bureau MA definitions for details.

policy or economic success. We first consider the employment/population ratio (EPR) over the respective periods. The EPR captures labor market tightness attributable to both unemployment and labor force participation, in which the latter includes factors such as older workers who lose their jobs and move onto disability (Autor et al., 2011). Favorable demand shocks will increase the EPR, suggesting that the original non-employed residents benefit. The EPR is approximated by dividing the number of employed residents, using US Bureau of Labor Statistics Local Area Unemployment Statistics data, by county population from the US Census Bureau.

Our second dependent variable is the poverty rate. Because place-based policy often targets poor areas, policy success in this regard requires that positive economic shocks reduce the poverty rate. This stands in contrast to sorting arguments that suggest that impoverished workers with less inclination to work self-sort into poor places and would receive few benefits from place-based policies. The 1990 and 2000 poverty rates are from the 1990 and 2000 Census of Population, while we use US Census Bureau SAIPE estimates for other years.[11]

We also use population growth as a dependent variable. Changes in the EPR suggest smaller (larger) migration adjustments as the new jobs are filled by previously unemployed and non-labor force participants. Because population data come from another source, the population models serve as confirmatory analysis of our EPR results.

Our regression models closely follow Partridge et al. (2012), though we consider different time periods and lagged adjustment effects to assess persistence. Variable details and sources of the explanatory variables can be found in the earlier paper. For each sub-sample, our base specification for a given county *i* located in state *s* is:

$$\mathbf{OUTCOME}_{\mathbf{is}(t)} = \alpha + \boldsymbol{\lambda}\mathbf{OUTCOME}_{is0} + \boldsymbol{\theta}\mathbf{ECON}_{is0} + \boldsymbol{\varphi}\ \mathbf{GEOG}_{is0} + \boldsymbol{\gamma}\mathbf{AMENITY}_{is} + \boldsymbol{\delta}\mathbf{DEMOG}_{is0} + \sigma_s + \varepsilon_{is(t-0)},$$

where the dependent variables are the EPR and the poverty rate measured in period *t* (e.g., 2000, 2010), and the population growth rate measured over the entire decade (1990-2000, 2000-2010).

[11]The 1990 and 2000 poverty rates are for the previous years, as the Census measures income in the prior year. The SAIPE estimates are poverty rates from a computer model. While the Census is more accurate, the correlation between the SAIPE and Census poverty estimates is approximately 0.95, suggesting that SAIPE data are reliable.

OUTCOME$_{is0}$ is the initial-period level of the dependent variable (except in the population model). A larger **λ** indicates greater persistence. **ECON** reflects economic activity, **GEOG** is a vector of variables that measure the location's access to the urban hierarchy, **AMENITY** contains measures of natural amenities, and **DEMOG** contains demographic/human capital attributes. The regression coefficients are α, **λ**, **θ**, **φ**, **γ**, and **δ**; σ_s are state fixed effects that account for common features within a state; and ε is the error term, assumed to be spatially correlated.[12]

The primary **ECON** variable is the industry mix employment growth for each period. It is the 'share' variable from shift-share analysis (Bartik, 1991; Blanchard and Katz, 1992) and is constructed by summing the products of the initial-period county industry shares and the national industry growth rates.[13] Industry mix employment growth represents the overall growth rate that would occur in a county if all of its industries grew at their respective national rates. Variation in county industry mix job growth is due solely to their having different initial industry composition. If an industry experiences a national or international demand shock, it influences the county's industry mix growth rate through its intensity in the county. A key advantage of our industry mix variable compared to past research, including Partridge et al. (2012), is that we use four-digit industry data versus one- or two-digit data—which provides a finer depiction of industry shocks.[14]

As migrants are attracted to a region with favorable (industry mix) demand shocks, there will be increasingly smaller increases in the EPR and smaller decreases in the poverty rate. The industry mix variable will be more strongly related to the EPR and poverty to the extent that local labor supply satisfies local labor demand shocks.

GEOG contains measures of agglomeration economies including spatial distance measures that reflect proximity to urban areas differentiated by their tier in the hierarchy. First is distance to

[12]The residual is assumed to be spatially correlated with neighboring counties in their Bureau of Economic Analysis functional economic region but independent of county residuals outside the region.

[13]Industry mix employment growth for a county equals $\sum_i(e_i/E)*gn_i$, in which e_i is county employment in industry i, E is total county employment, and gn_i is the national employment growth rate for industry i. Driven by national or international shocks, the industry mix growth rate is often used as an exogenous measure of total employment growth.

[14]The source of the four-digit industry level data is the EMSI consulting company. Typically, public data sources do not report detailed industry data at the county level for confidentiality reasons. EMSI produces accurate data for industries that are not publicly disclosed through an algorithm that uses many government data sources. For details of their procedure, see Dorfman et al. (2011).

the nearest urban center of *any* size including micropolitan areas. For a county that is part of a MA, this distance is from the population-weighted center of the county to the population-weighted center of the MA. For a nonmetropolitan county, distance is measured from the county center to the center of the nearest urban area.[15]

Beyond the nearest urban center, we include the incremental distances to more populous higher-tiered urban centers: incremental distance in kilometers from the county to reach a MA of any size; and the incremental distances to reach MAs of at least 250,000, 500,000, and 1.5 million people.[16] The largest category generally reflects national and top-tier regional centers. Generally, more remote counties have less economic growth and lower wage and land prices (Partridge et al., 2008, 2009). The **GEOG** vector also includes the county's population, as well as the population of the nearest/actual urban center to account for net urbanization economies. Finally, the vector includes the county land area in square miles.

We account for natural amenities (**AMENITIES)** with a 1 to 7 scale provided by the US Department of Agriculture using measures of climate, proximity to water, topography, etc. Three indicator variables are added for close proximity (within 50kms) to the Atlantic Ocean, Pacific Ocean, and the Great Lakes. State fixed effects control for policy differences or other state-specific omitted influences. Thus, the other regression coefficients are interpreted as the average response for *within*-state changes in the explanatory variables.

The **DEMOG** vector includes several variables associated with human capital and mobility, all measured in the initial period. There are five variables measuring race or ethnicity; four variables measuring county educational attainment; percent of the population that is female; percent of the population that is married, and the percent with a work disability.[17]

4. Empirical Results

The descriptive statistics and the regression results for the key variables are presented in

[15]For a one-county urban center, the distance term is zero. The MA population is based on initial-year population.
[16]For a county already located in a MA or micropolitan area, the incremental value to reach a micropolitan area or MA (of any size) is zero. See Partridge et al. (2008) for more details of the incremental distances and maps that illustrate their construction.
[17]In models that use 1990 as the initial year, there are only four race and ethnicity measures due to data availability.

Tables 1 and 2 respectively, with the MA results in the upper panels and the nonmetropolitan results in the lower panels. In Table 2, columns 1-3 report the 1990-2000 results for the 2000 EPR, 2000 poverty rate, and 1990-2000 population growth; columns 4-6 report the corresponding results for the 2000-2010 period.

The coefficients on the lagged EPR range from, 0.66 to 0.79 in the 2000 model, and from 0.71 to 0.94 in the 2010 model (columns 1 and 4). They are also larger in the latter period for the poverty rate (columns 2 and 5). While the lagged coefficients may not indicate unit roots, they do suggest a very high level of persistence in the disequilibrium adjustment process to spatially asymmetric shocks.

The metropolitan coefficient on the 1990-2000 industry mix variable is not statistically different from zero and is small for nonmetropolitan areas in the 2000 EPR model, suggesting that almost all of the newly created jobs went to outside residents. Yet, in both samples there were significant anti-poverty effects associated with the newly created jobs, possibly through wage effects, or through poor residents moving to better jobs. The corresponding industry mix coefficients for the 2010 employment rate range from 0.11 to 0.30. Although shocks to local labor markets were quite persistent in both two decades, a key difference is that there is less interregional economic migration after 2000, suggesting migration played a smaller role in the adjustment process with more local residents obtaining employment.

We now further explore whether the Great Recession altered adjustment patterns by splitting the latter decade into 2000-2007 and 2007-2010 sub-samples. We then re-estimate the model using 2007 and 2010 EPR as the dependent variable and the employment industry mix variables measured over 2000-2007 and 2007-2010 (not shown).[18] The results are similar with the coefficient on the lagged dependent variables ranging from 0.86 to 1.0 (slightly more persistent than before) and the respective sub-period industry mix coefficients being quite close to the overall 2000-2010 results. Hence, the recession and expansion periods experienced

[18]With the exception of the industry mix variables that which correspond to the respective sample periods, we use the same explanatory variables in both the 2000-07 and 2007-10 models. We experimented with models that omit the demographic variables but the results were robust to those changes.

adjustment processes similar to the overall period.

Columns 4 and 6 report the corresponding results when using 1990-2000 and 2000-2010 percent population growth as the dependent variable. Population growth as the dependent variable precludes the use of the lagged population variable as a measure of persistence. The population growth rate results confirm the EPR and poverty results. There is almost a one-for-one MA migration response in the 1990s to industry mix employment growth and about a 0.75 response in nonmetropolitan areas, whereas, the respective migration responses were both about 0.18 in the 2000-10 period, indicating smaller migration responses to economic shocks and larger responses in the EPR and poverty rate.

5. Discussion and Conclusion

The empirical literature reviewed above suggests that the spatial equilibrium view of regional economies is an imperfect representation of reality. Interregional convergence either does not always occur, or if it does, the required time lag is long—perhaps intergenerational. Our update of the empirical results for the US supports this perspective in showing that employment-to-population ratios and poverty rates are affected by demand shocks over the period of a decade; there is also a high degree of persistence in these measures of regional economic outcomes. Likewise, our empirical evidence suggests economic-based migration is declining in the US. Given this evidence, should place-specific economic policies be used to address high local poverty and/or low employment rates?

A key feature of the spatial equilibrium model is the high degree of mobility of resources. Demand or supply shocks that have differential impacts across regions will induce the flow of resources to areas of higher productivity (controlling for relative amenity appeal) increasing aggregate economic growth. In this process some regions gain, some lose, but mobile individuals and firms responding to incentives in expectation of higher levels of utility/profits are better off. Aggregate economic growth is increased. Place-targeted policy interventions that inhibit the required spatial mobility may trap individuals and firms in uneconomic regions or industries, sometimes at great public expense, both directly through taxes and also through sacrificing higher rates of economic growth.

But if resources are not mobile, what are the policy options?[19] One set of policies relates to increasing the mobility of resources. People-based policies such as health and wellness improvements, education and training, job search incentives, subsidized moving expenses and information and recruitment campaigns may increase mobility. The appropriate regulatory framework and removal of trade barriers between regions, for example, can facilitate the movement of financial capital. Natural resources, the ultimate example of immobile resources, may nevertheless be mobile among uses.

There may, however, be instances where it is too costly or impractical to achieve the mobility that would be required for the poverty reduction or employment improvement goals. Even if transfer payments were replaced by subsidies to encourage mobility, durable housing can cause a limited population response to a negative labor demand shock (Notowidigdo, 2011). Similarly, if the poverty and unemployment is concentrated among the elderly, people-based policies to improve labor mobility will not be appropriate. If education/skills, institutional, cultural or language gaps are very large, the cost of closing those gaps may be very high. Substantially 'emptying out' of regions will also preclude potential future development.

Beyond cost considerations, people-based policies to increase mobility may simply be ineffective. Financial incentives aside, historical, cultural and language barriers may inhibit mobility in a fundamental way, at least for current generations. Pockets of Aboriginal populations, or immigrant populations, may not be very responsive to economic incentives designed to improve mobility. Social, historical and cultural ties may translate into very high social and personal costs. Likewise, programs to improve skills may be ineffective, especially among older workers will also reduce people-based program effectiveness.

When conventional people-based policies designed to increase mobility fail to address

[19]This paper is silent on the mechanics of policy interventions, specifically whether they take the form of conditional transfers to regional/local governments or constitute direct involvement of the senior government. The considerations and criteria remain the same.

persistently lagging regional outcomes, place-based policies may be considered in addition to, though probably not replacing, people-based policies. Place-based policies, as defined here, have two basic characteristics. First, they are policies on the part of a senior level of government for a lagging region in its jurisdiction. Local policies are always ‘place-based’ and are therefore not at issue. This is consistent with the original distinction between targeting ‘place-prosperity’ versus ‘people-prosperity’ described by Winnick (1966) and revisited by Bolton (1992). While improving the well-being of ‘people’ in the region is the ultimate target of both, targeting places may be chosen as a means to this end.

A second characteristic of place-based policies is that they result in immobile local investments such that the local population will benefit from the policy only in the region. In this way, place-based policy introduces a barrier to mobility. Examples include physical infrastructure and support for businesses in a particular place. The geographic immobility of the object(s) of the policy is a key characteristic to avoid outcomes such as ‘brain drain’ where individual recipients of people-based policies may leave the region (and be better off), though the region is worse off (Artz et al., 2009; Beckstead et al. 2008).

The appraisal of whether policies for lagging regions should include site-specific investments or incentives, that is, place-based policies, will include two main considerations. One, are regional employment and poverty outcomes responsive to local demand shocks? Two, will long term benefits outweigh the costs? If increases in local jobs lead to employment of the local unemployed (rather than in-migrants or commuters) and poverty reduction, then local job stimulus may be warranted. Well-designed econometric analyses such as that conducted in this paper will answer this question. The magnitude of the response will suggest the attractiveness of place-based investments.

The second criterion, comparison of the benefits and costs of place-targeted interventions, is more complex. Full opportunity cost accounting is called for as resources are shifted from (presumably) higher productivity regions or cities. Both costs and benefits are likely to be distributed over decades. Current policy decisions leading to significant public regional investments will both create new opportunities and preclude others. Stimulus for developing a new industry may result in a long-term dependency (the infant industry problem). Certainly, place-specific investment will aggravate existing labor immobility. Transparency and monitoring are necessary components for public accountability and responsibility.

If place-based policies in the form of local/regional investments to create local jobs do not qualify by one or both of the above criteria, simple income transfers may be the most efficient way to address localized poverty or lagging employment. The evidence reviewed in this study and discussion suggests labor mobility is limited for reasons beyond what might be generated by income transfer payments. The costs of transfer payments arising from limited labor mobility also should be considered in setting national policies, such as those related to international trade (Autor et al., 2011). Reliance on household transfer payments alone is not an attractive long-term economic and political solution, however, and other options will likely need to be pursued, though they should not be at the expense of income transfer payments.

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Table 1: Descriptive Statistics

Variable	Obs.	Mean	Std. Dev.	Min.	Max.
Metropolitan Areas					
Industry mix emp. growth 1990-2000	1053	0.1680	0.0536	-0.2093	0.3545
Industry mix emp. growth 2000-2010	1040	0.0355	0.0478	-0.2118	0.2580
Employment population ratio 1990	1053	0.4678	0.0594	0.1215	0.7602
Employment population ratio 2000	1053	0.4845	0.0561	0.1537	0.6761
Employment population ratio 2010	1053	0.4479	0.0529	0.1494	0.6104
Poverty rate 1990[a]	1053	13.2675	6.2605	2.1802	56.8443
Poverty rate 2000[a]	1053	11.5538	5.1930	2.1174	35.8708
Poverty rate 2000[b]	1053	10.8815	4.4356	1.7000	31.7000
Poverty rate 2010[b]	1053	14.7118	5.1940	3.5000	35.8000
Population growth rate 1990-2000	1053	0.1807	0.1801	-0.1231	1.9207
Population growth rate 2000-2010	1053	0.1210	0.1490	-0.4529	1.1196
Non-Metropolitan Areas					
Industry mix emp. growth 1990-2000	1971	0.1297	0.0469	-0.0848	0.3511
Industry mix emp. growth 2000-2010	1963	0.0229	0.0548	-0.2308	0.3398
Employment population ratio 1990	1971	0.4323	0.0578	0.1947	0.8440
Employment population ratio 2000	1971	0.4554	0.0632	0.1899	0.8078
Employment population ratio 2010	1971	0.4429	0.0792	0.1827	0.8374
Poverty rate 1990[a]	1971	18.5307	7.9980	2.4017	63.1177
Poverty rate 2000[a]	1971	15.4994	6.6159	2.9252	52.3189
Poverty rate 2000[b]	1971	14.6075	5.6426	2.7000	42.2000
Poverty rate 2010[b]	1971	17.9166	6.3583	3.2000	49.1000
Population growth rate 1990-2000	1971	0.0740	0.1340	-0.2718	0.8819
Population growth rate 2000-2010	1971	0.0170	0.1000	-0.3800	0.8984

Notes: [a]Poverty rate data from the Census of population; [b]Poverty rate data from US Census Bureau SAIPE.

Table 2: Empirical Results: 1990-2000 and 2000-2010

	1990-2000			2000-2010		
	2000 Emp. Ratio	2000 Poverty Rt.	1990-2000 Pop. Chg.	2010 Emp. Ratio	2010 Poverty Rt.	2000-2010 Pop. Chg.
Metropolitan Areas						
Lagged Employment Ratio	0.7895** (20.24)			0.7081** (21.33)		
Lagged Poverty Rate		0.5981** (15.65)			0.8955** (24.39)	
INDMIX Employment Growth	0.0249 (0.99)	-4.5816** (-3.26)	1.0257** (8.1)	0.1068** (4.75)	-6.5262** (-4.02)	0.1762** (1.67)
N	1053	1053	1053	1040	1040	1040
R^2	0.9065	0.9197	0.5769	0.8705	0.9103	0.4476
Non-Metropolitan Areas						
Lagged Employment Ratio	0.6567** (19.47)			0.9401** (22.43)		
Lagged Poverty Rate		0.5825** (25.55)			0.9328** (41.05)	
INDMIX Employment Growth	0.0394** (2.45)	-6.748** (-4.81)	0.7461** (8.63)	0.297** (11.92)	-6.3197** (-5.24)	0.1821** (3.16)
N	1971	1971	1971	1963	1963	1963
R^2	0.8646	0.8958	0.5444	0.769	0.9077	0.4608

Notes: For the 1990-2000 period, poverty data are from the decennial census; for 2000-2010, they are from SAIPE. Robust t-statistics from STATA cluster command are in parentheses. ** indicates significance at 5%. In all models, control variables include: distance to nearest or actual Urban Center; incremental distance to a MA; incremental distances to MA > 250,000, > 500,000, and > 1,500,000 population; county population 1990/2000; population of nearest or actual MA 1990/2000; county area (sq. miles); amenity dummy variable represented by a 1 to 7 scale (USDA); proximity (within 50kms) to the Atlantic Ocean, Pacific Ocean, and the Great Lakes; state fixed effects; demographic variables including five ethnicity shares (four for 1990); four education shares; %females; % married; and % with a work disability.