

Gottlieb, Paul

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Agricultural preservation, large-lot zoning, and real estate development in New Jersey, USA

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Agricultural preservation, large-lot zoning, and real estate development in New Jersey, USA

Paul D. Gottlieb
Associate Professor
Department of Agricultural, Food, and Resource Economics
Rutgers University
55 Dudley Road
New Brunswick, NJ 08901 USA
Phone: 00-1-848-932-9122
Mobile: 00-1-732-501-1712
Email: gottlieb@aesop.rutgers.edu

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ABSTRACT

In municipalities on the urban fringe in the United States, large-lot zoning is commonly thought of as a rural preservation tool. It could preserve rural landscapes by making residential lots too expensive for most homebuyers; or it could simply lead to estates that resemble farms or retain significant forest cover. If a lot size restriction is set too low, however, it could have the unintended consequence of increasing the amount of backyard per housing unit without significantly reducing the number of units sold. For this reason, large-lot zoning is sometimes accused of being a major cause of “urban sprawl,” even as its advocates proclaim it to be the remedy.

This article evaluates these conflicting claims using a panel dataset on land cover and minimum-lot size zoning in 83 New Jersey municipalities between 1986 and 2002. It draws on a prior study that found the decline in homebuilding in this study area to be inelastic with respect to increases in municipal average lot size minima. That finding implies that residentially subdivided land area should rise with increased lot size minima, but of course not all of this land will be converted to suburban lawn cover. Three alternate regression specifications find no robust statistical relationship between municipal average lot size minimum and the ratio of residential to rural land cover. The conclusion is that large-lot zoning is an ineffective policy tool for reducing one common definition of “urban sprawl” at the municipal scale. A similar result is found with respect to easement purchases on farm and forestland within communities, which may not prove effective at reducing rural land conversion in the short run.

JEL classifications: R52 - Land Use and Other Regulations; R14 - Land Use Patterns; R31 - Housing Supply and Markets.

ERSA conference themes: G, M

1. Introduction

The loss of prime agricultural soils to development has long been regarded as a problem in both the United States (Raup, 1975; Lockeretz, 1986, 1989) and in Europe (Verzandvoort, 2009; Prokop, 2011). Policy responses to this problem vary, both between the two continents and within them. Planning approaches, permit reform, incentives, disincentives, and public acquisition are all on the menu of policy options for preserving prime farmland in these and other developed countries.

In contrast to the U.S., European governments conduct spatial planning at the national level. It should be remembered, however, that many European countries are no larger than individual U.S. states, which do have growth management plans and agricultural protection frameworks (DeGrove, 1992; Daniels and Nelson, 1986; AFT, 2005). U.S. states, especially in the Midwest and Northeast, are more likely than their European counterparts to lodge land use decision-making at the local level – although European municipalities are by no means powerless (Prokop, 2011). While Europeans have a reputation for very strong development restrictions compared to North Americans, recent progress reports for Europe suggest that urban sprawl is on the upswing and regulators are not meeting their rural preservation targets (Verzandvoort, 2009).

Perhaps the largest contrast between the two continents consists of culturally-mediated differences in regulatory or planning approach. Europeans have

experimented with nationwide rural impact fees charged to developers (Prokop, 2011). To a more limited extent, these programs exist in the U.S. in the form of local infrastructure fees or transferrable development rights. Equally common, however, are programs in which the state simply buys whatever rural properties or development rights it wishes to set aside (AFT, 2005). The ‘right to develop’ is generally assumed, so compensation in the U.S. tends to go from the taxpayer to the landowner, rather than the other way around.

Europeans appear to place greater emphasis than do Americans on the potential for brownfields development to relieve pressure on the countryside (Prokop, 2011). Middle-class Americans traditionally avoid city living (Jackson, 1987), so urban sprawl must be confronted and controlled almost entirely in a greenfield setting.

This last point of comparison leads to a discussion of what is perhaps the dominant rural preservation tool used in the United States: Minimum lot-size (MLS) zoning. This popular form of land use control does not restrict development directly. Instead, it places a floor on the size of the residential lot that surrounds each new home. The effect of such restrictions is to reduce suburban densities without forcing new homeowners back into the city.

Obviously, reducing residential densities is not the same as preserving agriculture on prime soils. That being said, if a lot size minimum is set at 5 acres (2 ha), then the entire residential lot is unlikely to be landscaped as a suburban lawn. If a lot size minimum is set at 20 acres (8 ha), then the parcel is likely to be farmed,¹ while a farm of greater than 20 acres will be a poor candidate for residential subdivision.

Lot size minima in the 5- to 20-acre range are common in the U.S. and 40 acres is not unheard of (Isakson, 2004; Newburn and Berck, 2011). These regulations protect local agricultural districts in the short run, although they are argued to impose significant wealth costs on large agricultural landowners (Shupe, 2004; Liu and Lynch, 2011). If such zoning restrictions were eventually relaxed, they could, in theory, allow higher density development to occur in concentric rings around a pre-existing town center. This type of zoning-implemented growth management scheme frequently takes place at the local level, so it has difficulty eliminating fragmentation at a larger level of geography, like the state or the county.

The goal of the present paper is to examine, empirically, the development and landscape effects of large-lot zoning across a sample of municipalities in the state of New Jersey, in the northeastern United States. Containing residential suburbs of both New York and Philadelphia, New Jersey is relatively affluent and is heavily settled at suburban densities. It is a “home rule” state that delegates significant land use powers to local governments. Its 566 municipalities have an average size of 15 square miles (39 km²), a fact that tends to inhibit the achievement of regional planning goals. For these reasons, the present paper will focus on the goal of rural land preservation entirely at the municipal level, where the key decisions are made. Forget about the state: Do municipal governments even achieve what they set out to accomplish with their comprehensive plans and agricultural zoning?

The dataset used for the present study has two main advantages for addressing this question: (1) It measures municipal lot size minima at three points in time,

allowing the use of panel techniques; (2) Its policy outcome measures utilize land cover data from aerial photography. This means that we can examine what policy makers really care about: the conversion of agricultural or wooded land to manicured lawns and impermeable surface (sealed soils). Challenges in the study include the need to aggregate zone-level data to the municipal scale, and a sample of only 83 communities, some of which were essentially built out at the start of the study period.

Many model specifications were run on the data. Although the signs on the coefficients matched the theory, estimates of the relationship between minimum lot size and increased residential land cover did not pass standard levels of statistical significance. The interesting policy conclusion is that while large-lot zoning may be effective at reducing the number of homes built in a municipality out to some horizon year (Gottlieb, *et al.*, 2012), it does not reduce the conversion of land from rural to residential cover for the simple reason that each new home is built on a larger lot. Large-lot zoning is therefore ineffective at curbing a peculiarly American form of urban sprawl: the conversion of farmland and forest to backyards, even as the number of housing units is reduced.

2. Prior literature on zoning and land conversion

Two theoretical articles are of interest. Using a model with a suburban housing sector and an agricultural sector, Moss (1977) concludes that minimum lot size requirements increase land prices, accelerate the process of rural land conversion, and increase total metropolitan land area. Using a residential economy model in a

semi-closed city with the rich living in the suburbs and the poor in the central city, Pasha (1996) concludes that suburban minimum lot size zoning flattens the rent gradient, reduces land prices, and leads to more residential land conversion and a larger metropolitan area. These works argue that the impact of minimum lot size zoning on land prices is an important mediating factor for determining its effect on land consumption. The fact that both authors predict an increase in residential land consumption with raw land prices moving in opposite directions, however, highlights the importance of assumptions in such models.

One can of course use data on land use or land use change as the dependent variable in a regression model of MLS zoning's effects. Using a logit model, Wu and Cho (2007) find that four types of local land use regulations *except* zoning ordinances significantly influence the probability of land development in five western states. Using expansion factors to proxy the acreage of each site, they estimate that land use regulations have reduced the amount of developed land by 10% throughout the whole region. Again, zoning ordinances were the exception to this finding on the development effects of local land use regulations.

Using county-level data from Ohio, Irwin, Hsieh and Libby (2001) find that the relative share of undeveloped to urban land is negatively and significantly influenced by zoning, represented by the proportion of land in each county that is formally zoned. Although this finding would appear to support the hypothesis that zoning facilitates land conversion, endogeneity is likely to be especially problematic in this case. In contrast to the choice of large versus small minimum lot size, the initial adoption of

zoning in a township that previously lacked zoning will clearly be driven by growth pressure and recent development.

Several U.S. studies have identified the expected relationship between large-lot zoning and lower residential densities, sometimes using subdivisions as the unit of analysis (McConnell, Wall, and Kopits, 2005; Lichtenberg and Hardie, 2007; Pendall, 1999). Because these studies measure housing unit density rather than landscape conversion, they are not strictly analogous to the present work. Their main value is that two out of three of these studies control for the possibility that zoning is endogenous (Pogodzinski and Sass, 1991). They therefore confirm that large-lot zoning is a binding regulation that reduces the number of supplied housing units while increasing the average size of each residential lot. These two findings will prove crucial to the mathematical framework presented below.

In summary, theoretical studies conducted at the metropolitan scale argue that minimum lot size zoning should lead to greater consumption of open land for residential use. Corresponding empirical evidence is thin to nonexistent, however, and studies have not been conducted at the scale most relevant to policy makers in the northeastern U.S. In the following section, we develop a simple conceptual framework for an empirical test of the impact of minimum lot size zoning on residential land conversion at a municipal scale. This is followed by a discussion of empirical setup, data, and results.

3. Conceptual framework

In his introductory text in natural resource economics, Field (2001) introduced a simple formula to describe the land consumption impact of minimum lot size zoning:

$$C = q(x) * als(x) \quad (1)$$

where C represents total land converted to residential use in a municipality over a given time period, $q(x)$ represents the number of new homes constructed as a function of minimum lot size x , and $als(x)$ represents the average lot size of these

new homes as a function of the minimum lot size: $als(x) = \frac{C}{q(x)}$. As constructed

here, the phrases “land consumption” and “lot size” could refer either to acres of legal ownership or to residential land cover after subdivision, provided that the same definitions are used for both C and for $als(x)$. To generalize the model further, x could be a weighted average of minimum lot sizes in a municipality that has more than one residential zone. Functions $q(x)$ and $als(x)$ both make sense at the municipal scale under a more aggregated interpretation of x .

We begin by defining $als(x)$ in terms of acreage of ownership units rather than land cover. We now add a strong, but defensible assumption, one that will be relaxed later. We assume that once the community selects x , all undeveloped land will be subdivided at the regulatory minima. Adding an assumption about proportionate development across different zones within the community (Gottlieb *et al.*, 2012) leads to a working assumption that $als(x) = x$. In other words, average future lot size in a municipality will be equal to average minimum lot size.

What could justify this strong assumption, since we know that MLS is a

one-sided constraint? The Tiebout framework of public finance suggests that incoming residents will prefer to buy lots that are no bigger than those of their neighbors in the same zone. Because local services in the U.S. tend to be financed by the property tax, this will ensure that new homeowners do not cross-subsidize the services of other newcomers (Tiebout, 1957; Hamilton, 1976). The only stable equilibrium for this game is one in which all new lots within a community are subdivided at the minimum legal size. Indeed, in a pure Tiebout world, this desire to prevent free-riding extends to incumbent homeowners as well; heterogeneity in the size of homes and backyards will occur across municipalities, never within them. For present purposes we assume uniformity of developed lots within residential zones but not entire communities, acknowledging the reality that income, housing, and zoning heterogeneity does exist within suburban municipalities.²

Using this assumption that $als(x) = x$, equation (1) can be rewritten with the municipal average minimum lot size x in place of average actual lot size in a community:

$$C = q(x)x \quad (2)$$

A community considering an increase in its average minimum lot size is concerned about the effect of an increase in x on total land consumption C . Take the first derivative of (2) with respect to x :

$$\frac{\partial C}{\partial x} = x \frac{\partial q(x)}{\partial x} + q(x) \times 1 \quad (3)$$

Multiplying through by ∂x :

$$\partial C = x \partial q(x) + q(x) \partial x \quad (4)$$

The expression on the right side of (4) is analogous to the economist's expression for the impact of a change in price on total revenue, given a price elasticity of demand. Equation (4) implies that change in total land consumed will be equal to current lot size times the change in the local demand for lots with increasing x (which the literature suggests is negative); plus existing lot demand times the change in the size of each lot (which the literature suggests is positive). As in revenue analysis, the change in C can be either positive or negative, which is what makes this an interesting empirical problem. And, as in revenue analysis, whether C rises or falls will depend on the elasticity of local housing demand with respect to average minimum lot size. The breakpoint, of course, is an elasticity with absolute value above or below 1.0.

We recently published an article on the effect of minimum lot size on municipal housing starts in this same study area, using federal building permit data (Gottlieb, *et al.*, 2012). Similar to the three density-oriented studies cited above, as well as others that address the effect of lot size zoning on the number of local housing units (Levine, 1999; Thorson, 1997; Glaeser, *et al.*, 2006; Newburn and Berck, 2011), this study finds a negative effect on equilibrium home production with increasing lot size minima, controlling for municipal zoned capacity and economic pull factors.

The coefficients on minimum lot size estimated in Gottlieb *et al.* (2012) can be converted into elasticities. For single-family housing units built in our 83 communities between 1996 and 2002, the elasticity with respect to a measure of municipal minimum lot size is -.35. By a very different pathway than in Moss (1977) and Pasha (1996), this empirical finding leads to a prediction that increasing

minimum lot size will lead to an increase in the total amount of land subdivided into residential lots in a community. By analogy to the economist's price-revenue problem, the fall in demand³ for local lots is inelastic, so the net change in total land consumption, ∂C in equation (4), must be greater than zero.

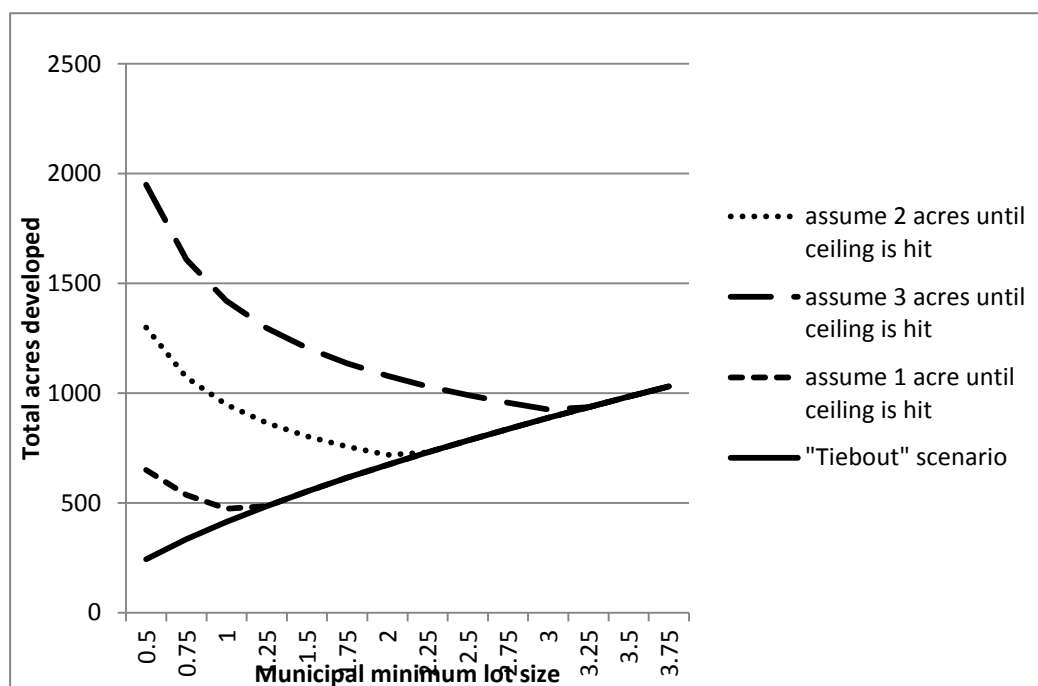
We now relax the assumption that new homes in a community are subdivided exactly at each zone's regulatory minimum. What alternative assumption can be made? Without Tiebout-Hamilton to guide us, there is little to predict how many affluent homeowners would be willing to tolerate lots so large that they might wind up cross-subsidizing the services of their neighbors.

An assumption that actual lots are subdivided at a fixed percentage above the regulatory minimum – e.g., that there is a systematic distribution of technically unconstrained lot sizes having a mean that rises along with the legal minimum – leads mathematically to the same predicted effect of zoning as a strict Tiebout assumption, but with higher values for $als(x)$ and C .⁴ An alternative assumption might begin with the observation that, if nothing else, the minimum lot size regulation must begin to bind once it gets large enough. Before that point, very small lot size minima could be ignored in favor of the uniformly large lots that developers and homebuyers seem to have preferred in this study area over the last few decades.

It is a simple matter to calculate and graph total municipal land consumption using an anticipated number of units to start with, the -.35 observed housing demand elasticity with respect to average minimum lot size, and alternative assumptions regarding the actual sizes of the lots subdivided under a one-sided regulatory

constraint. The main alternatives are shown in Figure 1. With actual lot sizes equal to the regulatory minimum, total land consumption increases steadily, as implied by the inelastic demand (see “Tiebout scenario” in the figure). When actual lot sizes are assumed fixed, say at 1 acre, total land consumption falls because there is falling demand with no increase in per-unit lot size; once the regulation begins to bind and lots are subdivided at the minimum, the “Tiebout” pattern of increasing land consumption takes over.⁵ The alternative scenarios incorporating some unconstrained behavior thus exhibit U-shaped residential land conversion as a function of minimum lot size, with the turning point occurring where the regulation begins to bind.

Figure 1. Total acres subdivided into private ownership units as a function of municipal minimum lot size

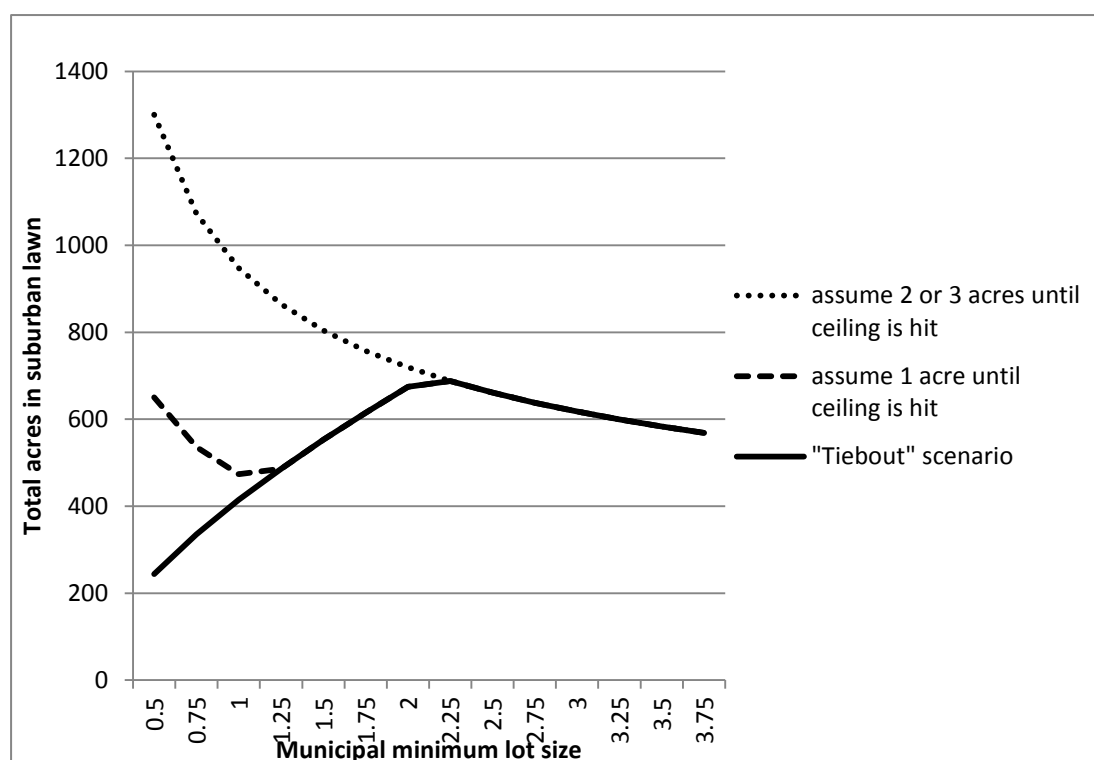


Starting units = 1,000. Elasticity of housing demand with respect to MLS = -0.35. Alternative scenarios represent actual lot sizes under a constraint that binds from below.

Figure 1 depicts total acres within residential ownership units. We now take the

housing unit demand underlying Figure 1 and use it to simulate changes in observed land cover, on the assumption that nobody will maintain a lawn larger than two acres, regardless of the size of the ownership unit. The results of this simulation are shown in Figure 2.

Figure 2. Acres converted to suburban lawn (maximum assumed two acres per lot) as a function of municipal minimum lot size



Starting units = 1,000. Elasticity of demand with respect to MLS = $-.35$. Alternative scenarios represent actual lot sizes owned, under a constraint that binds from below.

The phenomenon depicted in Figure 1 (total land conversion into private residential ownership units) is of interest; however, the data we have available to us is equivalent to that shown in Figure 2. We regard the bottom two scenarios of Figure 2 as more likely than the top scenario: new lots with a uniform size of two acres or greater are unlikely to exist in communities with MLS below 1.0. The two bottom scenarios in

Figure 2 suggest that the observed relationship between minimum lot size and land cover across the municipalities in our study area will exhibit a polynomial relationship that may or may not be detectable in the range MLS 1.0 to 5.0, given the inevitable noise in the data. We chose not to estimate a cubic function, but we did explore a quadratic. According to Figure 2, the most likely quadratic relationship for municipal land cover as a function of municipal MLS is concave, reflecting an eventual decline in unit demand with lawn cover per lot remaining constant, e.g., at two acres.

4. Empirical study design

4.1 Empirical specification

The data used to generate our dependent variable are the actual acres of land dedicated to different land use/land cover categories within each of 83 municipalities in 1986, 1995 and 2002. (This part of New Jersey became subject to greater regional planning control in the year 2004, so municipal-level analysis becomes less useful for subsequent years.) The land use share model is considered an ideal approach for characterizing aggregate land use in a jurisdiction (Miller and Plantinga, 1999). Because we are concerned with the amount of land converted from vacant to residential, we focus on residential and undeveloped land as our two land share categories.

As in Irwin, Hsieh and Libby (2001), define k as an index of categories of land use, where $k = a$ for undeveloped land and $k = b$ for residentially-developed land.⁶

With undeveloped land as the reference category, the expected share of residential and undeveloped land can be expressed using a multinomial logit model:

$$P_{ib} = \frac{\exp(X_i \beta_b)}{\sum_{k=1}^K \exp(X_i \beta_k)} \quad (5)$$

$$\text{and } P_{ia} = \frac{1}{\sum_{k=1}^K \exp(X_i \beta_k)} \quad (6)$$

where i indexes municipality, X are vectors of predictor variables, and β are vectors of estimable coefficients. Applying a logarithmic transformation to $\frac{P_{ib}}{P_{ia}}$

and substituting P_{ib} , P_{ia} with observed land use shares y_{ib} , y_{ia} ,⁷ the model is simplified as:

$$\ln\left(\frac{y_{ib}}{y_{ia}}\right) = X_i \beta_b + \varepsilon_i \quad (7)$$

where ε is an i.i.d. error term. The advantages of the logarithmic land share specification are that it is derived from a theoretically-grounded discrete choice model for the use of each acre, and it is not bounded by zero or one.

For convenience, define the dependent variable in (7) to be $\ln(r_i)$ for each municipality i . Adding the longitudinal dimension of our data, a fixed effects model is specified as:

$$\ln(r_{it}) = \alpha_i + \lambda_t + \beta X_{it} + \varepsilon_{it} \quad (8)$$

where t represents year, α represents a set of municipal-specific fixed effect dummies, λ is a dummy variable capturing time trend.

The fixed effects model is designed to control for sources of time-invariant

municipal heterogeneity, reducing omitted variable bias. Indeed, one can argue that unobserved heterogeneity will be especially important in a locally-diverse “Tiebout paradise” like New Jersey, with its 566 powerful towns.

4.2 *Data sources and variables*

Data have been collected for 83 municipalities in the Highlands region of New Jersey, with repeated observations for each municipality at years 1986, 1995 and 2002 (N=249). These municipalities are in the upland portions of Bergen, Passaic, Morris, Sussex, Warren, Somerset, and Hunterdon counties (see Appendix). All are therefore in the greater New York City metropolitan area. The land cover data come from New Jersey Department of Environmental Protection (NJDEP), which conducts aerial photography of the entire state at least once each decade. Maps based on 1:24,000 aerial photography are interpreted by state experts and digitized for further GIS analysis.⁸ The dataset classifies land cover using the comprehensive Anderson classification system (Anderson et al., 1976). For our study, residential land is defined as the sum of all residential categories; developable land is defined as the number of acres in agriculture or forest, along with a few open space categories in urban settings. Wetlands, barren land, and public recreation acres are not included in the undeveloped land category because they are not considered to be developable.

The 83 New Jersey Highlands communities sustained rapid urban growth during the study period, as has been documented elsewhere (Phelps and Hoppe, 2002). From 1986 to 2002, undeveloped/developable land within the study area declined by 39,300 acres (15,900 ha), accounting for about 5% of the region’s 775,800 acres

(314,000 ha), while residential land use increased by 28,500 acres/11,530 ha (this means, of course, that approximately 10,800 undeveloped acres converted to other improved categories, such as commercial or infrastructure). The extent of development varied quite a bit across the communities, as did zoning policies (see Table 1).

Historical zoning data for the Highlands region were collected under the terms of National Science Foundation Grant # SES 0523309. Each of the 83 municipalities had adopted several lot size requirements covering particular residential and agricultural zones in each year.⁹ To represent overall strictness of minimum lot size requirements at the municipal level, the weighted average minimum lot size (WMLS) was calculated for each municipality, with acres for each lot size category serving as the weights. This exercise was also conducted for the entire region, leading to the finding that the region as a whole changed from a weighted average minimum lot size of 2.2 acres in 1980 to 3.5 in 2004. This evidence of regulatory change is fortunate, because fixed effects analysis requires that the covariate of interest change over time for several places. Among the 83 communities in the sample, 79 had a different minimum lot size in place at the end of the study period than they had at the beginning.

We collected a set of independent variables that are both fixed and time-varying, even though the former are not used in fixed effects models. BOROUGH, a dummy variable, indicates whether the municipality is a borough/town as opposed to a township. Constituting the state's historical market centers, boroughs and towns in

New Jersey tend to be older, denser, and much closer to build-out than townships. This could affect the relationship between zoning and observed development (see below). Data were collected on distance to New York City (NYCDIST), number of highways (HIWAYS) the existence of any highway (ANYHIWAY), and the existence of a commuter rail station (RAIL) expressed as a dummy.

Time dummies were constructed for 1995 and 2002, with 1986 the omitted category (DUM1995; DUM2002). Several time-varying independent variables were collected to help explain residential land share in each municipality. These are median household income (MHINC),¹⁰ school general fund budget per pupil (PPEXP), crime rate per 1000 people (CRIME), and the interaction terms NYCDIST*DUM1995 and NYCDIST*DUM2002. Although NYCDIST itself is time-invariant, a fixed effects model can use interaction terms to explore changes in the impact of a fixed variable over time (Wooldridge, 2002). NYCDIST*DUM1995 and NYCDIST*DUM2002 are designed to capture changes in the density/development gradient within the New York metropolitan area.

In a review article, Quigley and Rosenthal (2005) highlight the need to understand the ways that multiple growth management tools work together. Latent factors such as differential enforcement and staffing at the local level may also be important, these authors argue. We maintain that the fixed effects approach effectively handles omitted covariates related to local administrative and political attitudes, which should be reasonably persistent in our study period. One policy variable, however, deserves special attention, and that is the legal protection of open space

through deed restriction or outright purchase.¹¹

In the early 1990s, the State of New Jersey passed some of the largest open space referenda ever seen in the U.S., with bond funds dedicated to the purchase of parks, nature preserves, and development easements on working farms (GSPT, 2008). The magnitude of these purchases has varied widely across New Jersey's municipalities and across time, with local political preferences, the financial situation of individual farmers, and state conservation policies all playing a role in their geographic distribution. We have collected data on the number of acres of legally-protected open space by municipality for each year.¹² These acres are a subset of the farm, forest, wetland, and barren acres measured in the Department of Environmental Protection's land use/land cover database. They should, in theory, include all of the acres identified in that database as recreational.

Using this official open space inventory, we build a time-varying covariate, OSPCT, which is defined as legally-protected open space acres as a percentage of total municipal land area in 1986/1995/2002.¹³ To the extent these protected open space acres are included in the land cover definition of open space that is used to calculate land share ratio r_{it} , OSPCT should have a negative coefficient in our regressions. This is because the more farm and forest acres that appear developable on the basis of aerial photography, but which are actually barred from development on legal grounds, the less residential development we should see, other things equal. On the other hand, farm and forest acres that are permanently preserved also have a likely "development magnet" effect (Roe, Irwin, and Morrow-Jones, 2004) that could

offset the negative coefficient we would otherwise expect on OSPCT. The net effect of OSPCT is indeterminate.

5. Empirical results

The fixed effects model shown in (8) above was run on the 83 municipalities for the three study years (N=249) (see Table 2 for results). Four strong outliers were identified.¹⁴ For now, these four municipalities were removed from the dataset, keeping the panel balanced but reducing sample size from 249 to 237.

Approximately half of the sample consists of “boroughs” or “towns” that were largely built out before zoning was invented. In these places, minimum lot size zoning does not necessarily bear a causal relationship to subsequent density or development. The cross-sectional relationship between zoning and residential lots in these municipalities could include widespread nonconforming uses that have been ‘grandfathered.’ For this reason, fixed effects models were also run with the boroughs and towns removed. These results are labeled “Townships only” in Tables 2 and 3.

Finally, whether using full or partial samples of the 83 communities, the fixed effects model was run with weighted minimum lot size specified both linearly and as a quadratic, as suggested by Figure 2.

The estimated coefficients on WMLS or WMLS², as with all zoning covariates, could potentially be subject to endogeneity bias. Using exogenous variables in the model along with some additional instruments, we implement the Spencer-Berk

variant of the Hausman test, and find that the fixed effects model exhibits no significant correlation between minimum lot size and the error term.¹⁵

Before concluding that the coefficients on the zoning variables are unbiased, as in previous studies that rely on the Hausman test, we address possible failures of fixed effect assumptions that relate to time (so-called strict exogeneity). In the present case, the fixed effects model assumes that there is no feedback between residential development shocks in time t and lot size regulations enacted in time $t+1$, a situation that would create a correlation between one of the errors and one of the independent variables. But in fact, such a correlation across time seems likely on behavioral grounds. Voters are likely respond to recent growth in the amount of developed land with tighter lot size controls, creating precisely this feedback effect.

The solution to this problem is to estimate the model with a weaker set of restrictions, known as sequential moment restrictions (Chamberlain, 1992). We first need to see if the problem exists. This can be done by limiting the observations in the fixed effects model to 1986 and 1995 and regressing the dependent variable on leading, instead of lagging, independent variables. Statistically significant coefficients on the leading covariates indicate potential feedback problems (Wooldridge, 2002). When this was done on our 1986 and 1995 data, none of the leading variables had estimated coefficients significantly different from zero at the 5% confidence level, and WMLS was not even close to this cutoff. Strict exogeneity appears to hold in this panel dataset. We also ran diagnostic tests for heteroskedasticity and spatial autocorrelation in the residuals: neither were found to be a problem.

The fixed effects models reported in Table 2 show a pattern of results that continues to hold throughout several alternative regression specifications. Simply put, the signs on the estimated coefficients for WMLS and WMLS² are exactly as predicted under the Tiebout scenario of Figure 2, but levels of statistical significance are either borderline or clearly insignificant. (We effectively use the more conservative one-tailed t-tests.) A comparison of significance levels suggests that the concave quadratic form for WMLS may be preferable to the linear form. Overall, however, the results suggest that the fixed effects test of the impact of WMLS on land consumption using a sample of 80 or 40 communities may lack statistical power.

It should be noted that the regression results with and without the four outliers produced similar results with one exception. The township-only model showed strong statistical significance for the quadratic coefficients on WMLS and WMLS² when Rockaway Township was included in the model, but not when it was omitted. It turns out that Rockaway Township is home to a military facility that is undevelopable, but which is nevertheless subject to Rockaway's 20-acre minimum lot size zoning. Municipal-wide WMLS for Rockaway is at the upper end of the distribution for our sample. At the same time, this township experienced less change in land cover for reasons that are similar to, but are not formally reflected in, the variable OSPCT. It follows that the regression results with Rockaway included are not robust or generalizable. Rockaway, a high-leverage observation, is omitted from Table 2's regressions along with three other strong outliers.

5.1 Alternative models

Several alternative specifications were run on the data, with similar results. The alternative models tested included some with lag structures. Two models are worthy of note. We ran the following first-differences model, which is roughly equivalent to fixed effects (see, e.g., Wooldridge, 2002, pp. 284-285):

$$\ln(r_{it}) - \ln(r_{i,t-1}) = \gamma_{95,02} + \beta(X_{it} - X_{t-1}) + e_{i,t,t-1} \quad (9)$$

where $t = 1995, 2002$ and $\gamma_{95,02}$ is a dummy for the second span of years 1995 to 2002

In addition, we ran a model for marginal land consumption on undeveloped land, conditional on WMLS in the base year. For the year 1995 only, we had access to minimum lot size data for land that had not yet been developed. This model, specified in (10) below, is similar to that estimated in Gottlieb, *et al.* (2012), except that the dependent variable is the number of acres converted to residential use between 1995 and 2002. It differs from models (8) and (9) in that WMLS is measured over undeveloped land in each municipality in 1995, rather than over the entire municipality. We estimated a negative binomial count data model on the acres converted to residential use and used a two-step IV technique to rule out endogeneity on the zoning and open space variables (see Gottlieb, *et al.*, 2012 for details). In addition to focusing on greenfield development, this is the only model of the three that is able to incorporate time-invariant covariates. Because it does not use a ratio r_i as its dependent variable, it must also include a measure of developable acres or municipal land area as a control variable within the X vector:

$$da_{i,2002} - da_{i,1995} = f(\beta X_{i,1995}) + \epsilon_i \quad (10)$$

where $da_{i,t}$ = residentially developed acres in municipality i in time t

Table 3 reports regression results for the first differences model shown in (9).

As expected, results for the two policy variables are similar to those from the fixed effects models of Table 2. Table 4 reports regression results for the model shown in (10). Coefficient signs on WMLS are once again as predicted, but some p-values now fall below 10%, even with the older boroughs and towns included. A model with only the 40 townships was not run because it would have too few degrees of freedom; instead, borough status was represented by a dummy variable.

Table 4 highlights the potential for achieving more precision in estimating the hypothesized policy relationships by looking at marginal development on undeveloped land only. Unfortunately, 1995 is the only year for which we currently have MLS data on undeveloped land, as opposed to entire municipalities.

5.2 Results on the permanent open space variable

In land share ratio models that count legally-protected open space within the dependent variable as ‘undeveloped,’ the coefficient on OSPCT is not statistically significant (Tables 2 and 3). The same is true for the marginal model of Table 4. We interpret this to mean that the negative effect on residential development of the legal restriction signified by OSPCT is offset by its tendency to attract more development to those acres that are truly developable. This interpretation can be confirmed by excluding legally-protected acres from the definition of developable land when calculating $\ln(r_i)$. This alternate model (not shown) generated a positive coefficient on OSPCT, presumably because the development magnet effect is all that

remains when the dependent variable is redefined in this way; only land that is truly developable is relevant. Attempts to preserve open space in a municipality by outright purchase may therefore be self-defeating, as observed elsewhere (Roe, Irwin, and Morrow-Jones, 2004).

6. Conclusion

The findings on the relationship between MLS and land consumption in this dataset are inconclusive, but they suggest that we are on the right track. Although the present tests of this relationship may lack statistical power, the hypothesis of a concave relationship (Figure 2) remains more reasonable than the alternatives.

The current failure to reject the null hypothesis on WMLS, however, has policy significance when combined with the theory presented here. We may be seeing a relatively flat quadratic relationship between WMLS and land cover, that, given the noise in the data, is hard to distinguish from a horizontal line. If a community thought it could reduce the conversion of its land from farm and forest to suburban lawns by increasing average minimum lot size, it would be sorely mistaken. A large-lot policy appears to be ineffective at achieving this goal. The same may be said of open-space set-asides, at least in the short run.

An important next step in this research is to develop a spatial cross-tabulation of acres by development/landscape status and zoning category, in a dataset with a time dimension. This would enable researchers to drop the weighted-average MLS measure for municipalities and examine residential development across different

minimum lot size categories directly. A happy side effect would be more data at the high end of MLS, for the simple reason that these extremes are effectively averaged away in the present study. “Marginal” greenfield analysis for the 1986-1995 and post-2002 periods could also be conducted, leading to a possible fixed effects analysis of marginal growth on greenfield land in each base year.

At present, we have failed to identify a robust relationship between a municipality’s weighted average minimum lot size on all of its residential territory, and the proportion of its developable land that is in some form of residential land cover. This is true even for a township-only sample, where almost all of the zoning and development will have occurred after World War II. These findings are policy relevant. In addition, we have established some theoretical benchmarks for continued analysis of the relationship between municipal lot size minima and the amount of land that is subdivided into residential lots, or is converted to suburban land cover, such as residential lawns.<end>

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¹ Most U.S. states offer real estate tax breaks to parcels above a certain size (e.g., 5 acres in New Jersey) that are engaged in agricultural production. These tax provisions also incentivize

agricultural activity on large parcels.

² The Tiebout-Hamilton argument is really about homogeneous housing values, not about equal lot sizes. Many communities, however, enact additional zoning restrictions that, either directly or indirectly, specify the ratio of the size of the structure to the size of the lot. Such ordinances, which include regulations on bulk, setback, side yards, and floor area ratio, tighten the already strong relationship between a home's assessed value and its lot size. A summary of the empirical evidence on Tiebout-sorting may be found in Dowding, *et al.* (1994).

³ Because short run housing supply is regarded to be highly elastic, the phrase "demand elasticity" will be used here even though the dependent variable used in Gottlieb *et al.* (2012) is actually a measure of equilibrium home production.

⁴ One may verify this effect by specifying $als(x)=kx$ where $k > 1$. Equation (4) will now be the same above, except multiplied through by the constant k .

⁵ Note that changing assumptions about the sizes of the lots actually developed does not alter the fact that our prior study (Gottlieb, *et al.*, 2012) estimated a uniformly negative coefficient on single family building permits as a function of regulatory lot size minima -- *not* as a function of actual lot sizes, which were unobserved. Figures 1 and 2 below must therefore be constructed in terms of municipal-level minimum lot size, which also happens to be the policy variable of interest. Actual lot sizes, being unobserved, are a matter for simulation and sensitivity testing in these figures. No attempt is made here to revisit the theory underlying Gottlieb *et al.* (2012)'s empirical finding on local housing demand in light of the figures' temporary relaxation of the strict Tiebout assumption.

⁶ The analysis of land shares is relative: categories a and b need not comprise 100% of municipal land area (Miller and Plantinga, 1999). From 1986 to 2002, the total acres included here for each town may actually change due to the commercial and industrial development of vacant land, which is excluded from the analysis. This exclusion is justifiable on the grounds that (1) commercial development is not part of our theoretical model, (2) commercial development is a relatively small proportion of the land area of these suburban towns (see Table 1), and (3) retail development, which is a large share of business acres in most towns, tends to follow residents. Further analysis confirms that a focus on residential development is equivalent to a focus on all development. When category b was defined to include all developed acres, so that the sum of categories a and b completely exhausted municipal land area in each year, results were virtually identical to the model of residential development presented in Table 2. These results for the ratio of total developed to undeveloped land in the Highlands townships are available on request.

⁷ One may use either percentages or raw acres for y_{ib} and y_{ia} , since the denominator of any share calculation, total municipal land area, cancels out of the ratio.

⁸ We are grateful to Stephen Karp of the New Jersey Office of Smartgrowth for providing these GIS data in tabular form with categorized acreage by municipality.

⁹ Virtually all agricultural land in New Jersey is zoned residential with a minimum lot size, even if it is labeled an "agricultural preservation district" in the town's master plan. This practice reflects those development options maintained by farmers under law.

¹⁰ Given the study years, some income data had to be interpolated or extrapolated from data

reported in the decennial census.

¹¹ Transfer of development rights programs were not a significant factor in northwestern New Jersey at this time. Local urban growth boundaries are not authorized, and the State Development and Redevelopment Plan — a “tier” growth management system adopted in 1991 and governing development in this region until 2004 — is voluntary. Data on the local authorization or use of cluster options within MLS zoning regimes was not available.

¹² Our data sources allow us to identify acres in state and local farmland preservation programs, nonprofit land ownership, the state’s Green Acres program, and all state, municipal, and county parks and preserves. The one gap in the data are conservation easements owned by nonprofit organizations.

¹³ In this study, recreation acres (typically city parks) are excluded from both the legal open space inventory OSPCT and from the measure of developable land used to create the dependent variable. These acres are clearly distinguishable from the air and may be considered to be “developed/civic infrastructure.”

¹⁴ All of these four observations have studentized residuals exceeding ± 3 , the Cook’s D above $4/n$, and the DFITS greater than $\pm 2\sqrt{k/n}$. They also stand out uniquely in the residual plot over the predicted values and the normal Q-Q plot. These criteria for outliers are used throughout the analysis.

¹⁵ See Spencer and Berk (1981). This variant of the Hausman test is required because minimum lot size is not the only variable we expect to be endogenous, it is simply the one whose endogeneity we wish to test. We assume that MHINC, OSPCT, and PPEXP are also endogenous, and we jointly test MLS and MLS^2 for endogeneity by comparing one IV model to another (normal Hausman always compares IV to OLS). Percentage of land in steep slopes, with high conservation value, with prime agricultural soils, and with rail access are used as municipal-level instruments. Results of these endogeneity tests are available on request.

Tables

Table 1a
Summary statistics for all Highlands municipalities by year (N=83)

Variable	Symbol	Mean	Min	Max	Source
Total municipal acres ⁽¹⁾		9023.87	86.38	47097.77	New Jersey LU/LC database
Proportion residential in 1986		0.24	0.04	0.75	New Jersey LU/LC database
Proportion residential in 1995		0.26	0.05	0.75	New Jersey LU/LC database
Proportion residential in 2002		0.28	0.06	0.76	New Jersey LU/LC database
Proportion developable in 1986		0.56	0.09	0.89	New Jersey LU/LC database
Proportion developable in 1995		0.54	0.06	0.88	New Jersey LU/LC database
Proportion developable in 2002		0.51	0.05	0.85	New Jersey LU/LC database
Proportion commercial/industrial in 1986		0.07	0.00	0.29	New Jersey LU/LC database
Proportion commercial/industrial in 1995		0.06	0.00	0.28	New Jersey LU/LC database
Proportion commercial/industrial in 2002		0.07	0.00	0.27	New Jersey LU/LC database
Proportion protected open space in 1986	OSPCT	0.09	0.00	0.49 [‡]	(2)
Proportion protected open space in 1995	OSPCT	0.11	0.00	0.55 [‡]	(2)
Proportion protected open space in 2002	OSPCT	0.16	0.00	0.66 [‡]	(2)
Ratio of residential to developable acres 1986	r_{ij}	0.77	0.06	8.05	See text
Ratio of residential to developable acres 1995	r_{ij}	0.89	0.07	12.49	See text
Ratio of residential to developable acres 2002	r_{ij}	1.00	0.08	14.85	See text
Borough dummy	BOROUGH	0.41	0.00	1.00	NJ Legislative District Data Book
Distance to New York City in miles	NYCDIST	39.54	21.90	62.61	Author map analysis
Railroad dummy	RAIL	0.20	0.00	1.00	New Jersey Transit
Per pupil expenditures 1986	PPEXP	\$4,759	\$3,386	\$7,215	NJ Legislative District Data Book
Per pupil expenditures 1995	PPEXP	\$8,430	\$6,153	\$12,818	NJ Legislative District Data Book
Per pupil expenditures 2002	PPEXP	\$10,340	\$7,750	\$14,154	NJ Legislative District Data Book
Median household income 1986	MHINC	\$62,928	\$34,350	\$126,582	US Census
Median household income 1995	MHINC	\$72,279	\$36,393	\$141,279	US Census
Median household income 2002	MHINC	\$75,899	\$37,758	\$141,948	US Census
Weighted average MLS 1986	WMLS	1.69	0.13	9.90	Original archival research
Weighted average MLS 1995	WMLS	1.78	0.13	9.52	Original archival research
Weighted average MLS 2002	WMLS	2.13	0.13	10.14	Original archival research
Squared WMLS 1986	WMLS ²	5.27	0.02	97.98	Original archival research
Squared WMLS 1995	WMLS ²	5.59	0.02	90.55	Original archival research
Squared WMLS 2002	WMLS ²	8.52	0.02	102.78	Original archival research
Crime rate in 1986 (per 1000)	CRIME	23.36	5.20	71.50	FBI Uniform Crime Reports
Crime rate in 1995 (per 1000)	CRIME	16.76	5.00	62.40	FBI Uniform Crime Reports
Crime rate in 2002 (per 1000)	CRIME	13.70	3.80	49.20	FBI Uniform Crime Reports

(1) Excludes water cover

(2) NJ State Agricultural Development Board; NJ DEP (Green Acres program); Municipal open space plans; Municipal open space inventories

Table 1b
 Summary statistics for Highlands townships only (N=42)

Variable	Symbol	Mean	Min	Max	Source
Total municipal acres ⁽¹⁾		15213.36	1791.31	47097.77	New Jersey LU/LC database
Proportion residential in 1986		0.15	0.04	0.43	New Jersey LU/LC database
Proportion residential in 1995		0.18	0.05	0.46	New Jersey LU/LC database
Proportion residential in 2002		0.20	0.06	0.47	New Jersey LU/LC database
Proportion developable in 1986		0.67	0.20	0.89	New Jersey LU/LC database
Proportion developable in 1995		0.65	0.18	0.88	New Jersey LU/LC database
Proportion developable in 2002		0.62	0.18	0.85	New Jersey LU/LC database
Proportion commercial/industrial in 1986		0.04	0.00	0.16	New Jersey LU/LC database
Proportion commercial/industrial in 1995		0.03	0.00	0.13	New Jersey LU/LC database
Proportion commercial/industrial in 2002		0.03	0.00	0.14	New Jersey LU/LC database
Proportion protected open space in 1986	OSPCT	0.12	0.00	0.41	(2)
Proportion protected open space in 1995	OSPCT	0.15	0.00	0.48	(2)
Proportion protected open space in 2002	OSPCT	0.22	0.04	0.66	(2)
Ratio of residential to developable acres 1986	r_{ij}	0.32	0.06	1.69	See text
Ratio of residential to developable acres 1995	r_{ij}	0.38	0.07	1.87	See text
Ratio of residential to developable acres 2002	r_{ij}	0.44	0.08	2.13	See text
Distance to New York City in miles	NYCDIST	41.32	21.90	61.96	Author map analysis
Railroad dummy	RAIL	0.14	0.00	1.00	New Jersey Transit
Per pupil expenditures 1986	PPEXP	\$4,665	\$3,386	\$7,215	NJ Legislative District Data Book
Per pupil expenditures 1995	PPEXP	\$8,246	\$6,153	\$10,758	NJ Legislative District Data Book
Per pupil expenditures 2002	PPEXP	\$10,065	\$7,750	\$14,154	NJ Legislative District Data Book
Median household income 1986	MHINC	\$66,856	\$43,530	\$117,491	US Census
Median household income 1995	MHINC	\$77,248	\$48,995	\$134,860	US Census
Median household income 2002	MHINC	\$81,481	\$51,687	\$140,185	US Census
Weighted average MLS 1986	WMLS	2.10	0.46	4.98	Original archival research
Weighted average MLS 1995	WMLS	2.27	0.43	4.98	Original archival research
Weighted average MLS 2002	WMLS	2.89	0.43	10.14	Original archival research
Squared WMLS 1986	WMLS ²	6.07	0.21	24.77	Original archival research
Squared WMLS 1995	WMLS ²	6.84	0.19	24.77	Original archival research
Squared WMLS 2002	WMLS ²	11.94	0.19	102.78	Original archival research
Crime rate in 1986 (per 1000)	CRIME	22.10	5.20	51.90	FBI Uniform Crime Reports
Crime rate in 1995 (per 1000)	CRIME	15.47	6.40	48.30	FBI Uniform Crime Reports
Crime rate in 2002 (per 1000)	CRIME	11.72	4.50	49.20	FBI Uniform Crime Reports

(1) Excludes water cover

(2) NJ State Agricultural Development Board; NJ DEP (Green Acres program); Municipal open space plans; Municipal open space inventories

Table 2

Determinants of the ratio of residential to undeveloped acres in New Jersey municipalities, 1986 to 2002

FIXED EFFECTS MODELS

	<i>Entire sample</i>		<i>Entire sample</i>		<i>Townships only</i>		<i>Townships only</i>	
	Parameter estimate	p-value	Parameter estimate	p-value	Parameter estimate	p-value	Parameter estimate	p-value
CONSTANT	-1.59392	<.0001	-1.57456	<.0001	-1.5734	<.0001	-1.52744	<.0001
PPEXP	-0.00000122	0.8958	-0.00000117	0.9004	0.00003814	0.008	0.00003544	0.0134
MHINC	0.00000604	<.0001	0.00000605	<.0001	0.00000433	0.0204	0.00000421	0.0249
CRIME	-0.00465	0.0011	-0.00473	0.0009	-0.00684	0.0004	-0.007	0.0003
OSPCT	-0.01147	0.9449	0.00781	0.9622	-0.27007	0.1562	-0.22346	0.237
WMLS	0.04462	0.1388	0.01905	0.1417	0.05363	0.0636	0.01606	0.2285
WMLS ²	-0.00367	0.3463			-0.00554	0.1413		
DUM1995	-0.14176	0.0407	-0.14514	0.0359	-0.27169	0.0032	-0.26227	0.0046
DUM2002	-0.11209	0.1758	-0.11259	0.1738	-0.30316	0.0132	-0.28211	0.0209
NYCDIST*DUM1995	0.00528	<.0001	0.00537	<.0001	0.00601	<.0001	0.00604	<.0001
NYCDIST*DUM2002	0.00715	<.0001	0.00712	<.0001	0.00868	<.0001	0.00846	<.0001
N*	3x79		3x79		3x40		3x40	
Adj R ²	0.99		0.99		0.99		0.99	
F test for joint significance of WMLS and WMLS ²	1.54	0.2185			1.86	0.1638		

* Four outliers were omitted from the entire sample and two from 'townships only'.

Table 3

Determinants of the ratio of residential to undeveloped acres in New Jersey municipalities, 1986 to 2002
 FIRST DIFFERENCE MODELS, TWO PERIODS (T=3)

	<i>Entire sample</i>		<i>Entire sample</i>		<i>Townships only</i>		<i>Townships only</i>	
	Parameter estimate	p-value	Parameter estimate	p-value	Parameter estimate	p-value	Parameter estimate	p-value
CONSTANT	0.16873	<.0001	0.03308	<.0001	0.09069	0.1055	0.09543	0.086
PPEXP	-0.0000921	0.1938	0.0000706	0.1547	0.0000227	0.0671	0.00002166	0.0775
MHINC	0.0000505	0.001	0.0000151	0.0007	0.0000053	0.009	0.00000526	0.0096
CRIME	-0.00153	0.1851	0.00114	0.2781	-0.00205	0.1968	-0.00216	0.1721
WMLS	0.03545	0.1477	0.00833	0.8216	0.02813	0.2983	0.00966	0.4373
WMLS ²	-0.0036	0.1448			-0.00273	0.4406		
DUMMY 1995/2002	-0.03894	0.0841	0.02241	0.0662	0.01048	0.7619	0.00764	0.8237
N*	2x81		2x81		2x40		2x40	
Adj R ²	0.15							
F test for joint significance of WMLS and WMLS ²	1.1	0.3357			0.6	0.55		

* Two outliers were omitted from the entire sample and two from 'townships only'.

Table 4

Determinants of acres converted to residential use in New Jersey municipalities, 1995 to 2002
 NEGATIVE BINOMIAL COUNT MODEL FOR ACRES CONVERTED TO SINGLE FAMILY

	<i>Entire sample</i>		<i>Entire sample</i>	
	Parameter estimate	p-value	Parameter estimate	p-value
CONSTANT	5.86221	0.000	5.87004	0.000
SPATIAL LAG OF DEPENDENT VAR	-0.00254	0.003	-0.00246	0.006
WMLS	0.32521	0.078	0.12928	0.071
WMLS ²	-0.02431	0.171		
OSPCT	-0.38375	0.624	-0.24950	0.754
DEVELOPABLE ACRES	0.00013	0.000	0.00013	0.000
ANY HIGHWAY (DUMMY)	0.21753	0.258	0.19096	0.322
VIOLENT CRIME RATE	-0.31812	0.022	-0.34874	0.008
NYCDIST	-0.03254	0.060	-0.02985	0.086
BOROUGH	-1.06360	0.001	-1.11578	0.000
% LAND IN STEEP SLOPES	-0.03183	0.258	-0.02924	0.311
MHINC	0.00001	0.379	0.00001	0.260
POP DENSITY	-0.00018	0.004	-0.00020	0.003
SPATIAL LAG OF EMPLOYMENT	-0.00006	0.912	-0.00004	0.945
RAIL	-0.25631	0.347	-0.29091	0.294
N	83		83	
Log pseudo likelihood	-444.05		-444.70	

Appendix. List of municipalities used in the study

Municipality Name	County Name
Mahwah Township	Bergen
Oakland Borough	Bergen
Alexandria Township	Hunterdon
Bethlehem Township	Hunterdon
Bloomsbury Borough	Hunterdon
Califon Borough	Hunterdon
Clinton Town	Hunterdon
Clinton Township	Hunterdon
Glen Gardner Borough	Hunterdon
Hampton Borough	Hunterdon
High Bridge Borough	Hunterdon
Holland Township	Hunterdon
Lebanon Borough	Hunterdon
Lebanon Township	Hunterdon
Milford Borough	Hunterdon
Tewksbury Township	Hunterdon
Union Township	Hunterdon
Boonton Town	Morris
Boonton Township	Morris
Butler Borough	Morris
Chester Borough	Morris
Chester Township	Morris
Denville Township	Morris
Dover Town	Morris
Hanover Township	Morris
Harding Township	Morris
Jefferson Township	Morris
Kinnelon Borough	Morris
Mendham Borough	Morris
Mendham Township	Morris
Mine Hill Township	Morris
Montville Township	Morris
Morris Township	Morris
Morris Plains Borough	Morris
Morristown Town	Morris
Mountain Borough	Morris
Mount Arlington Borough	Morris
Mount Olive Township	Morris
Netcong Borough	Morris
Parsippany-Troy Hills Township	Morris
Pequannock Township	Morris
Randolph Township	Morris
Riverdale Borough	Morris
Rockaway Borough	Morris
Rockaway Township	Morris
Roxbury Township	Morris
Victory Gardens Borough	Morris
Washington Township	Morris
Wharton Borough	Morris
Bloomington Borough	Passaic
Pompton Lakes Borough	Passaic
Ringwood Borough	Passaic
Wanaque Borough	Passaic
West Milford Township	Passaic
Bernardsville Borough	Somerset
Far Hills Borough	Somerset
Peapack-Gladstone Borough	Somerset
Byram Township	Sussex
Franklin Borough	Sussex
Hamburg Borough	Sussex
Hardyston Township	Sussex
Hopatcong Borough	Sussex
Ogdensburg Borough	Sussex
Sparta Township	Sussex
Stanhope Borough	Sussex
Vernon Township	Sussex
Allamuchy Township	Warren
Alpha Borough	Warren
Belvidere Town	Warren
Franklin Township	Warren
Greenwich Township	Warren
Hackettstown Town	Warren
Harmony Township	Warren
Independence Township	Warren
Liberty Township	Warren
Lopatcong Township	Warren
Mansfield Township	Warren
Oxford Township	Warren
Phillipsburg Town	Warren
Pohatcong Township	Warren
Washington Borough	Warren
Washington Township	Warren
White Township	Warren