Title

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Journal

Health & Place, 16(4)

ISSN

1353-8292

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Publication Date

2010-07-01

DOI

10.1016/j.healthplace.2010.02.005

Peer reviewed

Social Disparities in Health: Disproportionate toxicity proximity in minority communities over a decade

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February 10, 2010

Post-print. Published in Health & Place 16(4): 674-683

Word count: 6,575

Running Head: "Disproportionate toxicity proximity"

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Social Disparities in Health: Disproportionate toxicity proximity in minority communities over a decade

Abstract

This study employs latent trajectory models measuring the level of toxic waste over a decade in the cities of six highly populated, ethnically diverse, counties in southern California from 1990-2000 in 3,001 tracts. We find that tracts with 15% more Latinos are exposed to 84.3% more toxic waste than an average tract over this time period and tracts with 15% more Asians are exposed to 33.7% more toxic waste. Conversely, tracts with one standard deviation more residents with at least a bachelor's degree (15.5%) are exposed to 88.8% less toxic waste than an average tract. We also found that these effects were considerably weaker when using the raw pounds of toxic waste rather than the toxicity-weighted measure, suggesting that future research will want to account for the toxicity of the waste.

Keywords: environmental justice, neighborhoods, longitudinal, race/ethnicity.

Social Disparities in Health: Disproportionate toxicity proximity to hazardous exposures in minority communities over a decade

Substantial disparities in health status persist among minority populations, and the siting of toxic waste in low income communities with high proportions of minority residents is a significant Public Health problem facing the United States. In the United States, African-Americans, Latinos, American Indians/Alaska Natives, Asians, and Native Hawaiian or other Pacific Islanders bear a disproportionate burden of disease (Centers for Disease Control and Prevention 2004). There is increasing evidence that minority and low income populations are burdened with a disproportionate share of residential proximity to hazardous substances such as lead, PCBs, wood dust, and air pollutants (Agency for Toxic Substances and Disease Registry 2006; Centers for Disease Control and Prevention 2005) as well as toxic waste (Been 1995; Bolin et al. 2002; Downey 1998; Hite 2000; Hockman, and Morris 1998; Krieg 1995; Mohai, and Saha 2006; Pastor, Sadd, and Hipp 2001). This puts those living in such communities at risk for exposures that may be related to numerous diseases and disabilities.

The extensive literature studying possible disproportionate proximity among racial and ethnic minority populations to toxic waste largely employs cross-sectional designs and analyses (Anderton et al. 1994; Baden, and Coursey 2002; Been 1995; Bolin et al. 2002; Downey 1998; Hite 2000; Hockman, and Morris 1998; Krieg 1995; Mohai, and Saha 2006; Pastor, Sadd, and Morello-Frosch 2004; Sadd et al. 1999; Stretesky, and Hogan 1998). Furthermore, this literature focusing on the potential for possible disproportionate proximity to hazardous wastes based on race and socioeconomic status yields mixed findings: whereas several studies suggest a positive relationship between the proportion minority in a neighborhood (i.e., usually measured as tracts, a Census Bureau defined unit of approximately 4,000 persons) and the number of toxic waste sites (Been 1995; Bolin et al. 2002; Downey 1998; Hite 2000; Hockman, and Morris 1998; Krieg 1995; Mohai, and Saha 2006; Pastor et al. 2001), some studies have not detected such a relationship for African-Americans specifically (Anderton et al. 1994; Baden, and Coursey 2002; Sadd et al. 1999; Stretesky, and Hogan 1998) or minorities more generally (Bowen et al. 1995). Likewise, whereas some studies find a negative relationship between a neighborhood's economic resources and the presence of toxic sites (Bolin et al. 2002; Downey 1998; Hockman, and Morris 1998; Krieg 1995; Mohai, and Saha 2006), others have failed to find such a relationship (Baden, and Coursey 2002; Boer et al. 1997; Davidson, and Anderton 2000).

Cross-sectional studies have limited ability to ascertain the causes of such proximity to toxic exposures among minority populations given that they only provide a snapshot of the process (Northridge et al. 2003). Indeed, there are competing viewpoints as to why minorities' disproportionate proximity to toxic sites is frequently observed in cross-sectional studies. Initiatives at both the federal and community level address the burden of environmental injustice in the low income communities in which minorities reside. Despite these efforts, more research is warranted to understand the differential burden of environmental hazards in minority and low income communities, as well as the *determinants* of these existing environmental inequities (Northridge et al. 2003).

Explanations of disproportionate proximity

Although studies are generally limited to cross-sectional data, it is nonetheless the case that numerous theories have arisen to explain disproportionate proximity to toxic waste. One perspective argues that minority members may suffer disproportionate proximity to toxic waste due to their limited neighborhood options (Massey, and Denton 1987; Massey, Gross, and Shibuya 1994), relegating them to the highest toxicity neighborhoods. Beyond these more limited options, if such individuals (i.e., particularly immigrants with language barriers) are less aware of the health risks posed by such sites in these neighborhoods, they may be less averse to entering such neighborhoods.

A second perspective argues that toxic waste sites are disproportionately placed into neighborhoods that already have a high percentage of minority residents. This may be because such neighborhoods experience potential obstacles to residents' banding together to prevent the placement of such plants in their neighborhoods. One study explored this question explicitly with data from Los Angeles County and indeed found evidence that toxic sites were more likely to be placed in minority neighborhoods with high percentages of minority residents than in non-minority neighborhoods (Pastor et al. 2001).

The notion that some neighborhoods may lack the political willpower to resist such sitings suggests that the social disorganization model from criminology may be insightful for explaining this political tussle leading to toxic waste site placement. This model posits that neighborhoods with more economic resources, residential stability, and ethnic homogeneity have greater ability to collectively combat crime and social disorder when it appears (Sampson, and Groves 1989; Shaw, and McKay 1942). This model naturally extends to the question of toxic waste siting: neighborhoods with more poverty, residential instability and ethnic heterogeneity are likely least able to collectively resist the placement of such sites in their neighborhood, or to pressure owners to reduce the toxic waste emitted from existing plants (Hamilton 1993; Pastor et al. 2001). Building on these ideas, one study which examined neighborhoods in Los Angeles County found that neighborhoods with more *ethnic churning* (i.e., racial/ethnic transition) were *least* able to collectively resist the placement of such sites in their neighborhood (Pastor et al. 2001). It is possible that such neighborhoods also lack the ability to pressure existing plant operators to reduce the toxic waste emitted from existing plants.

Whereas the social disorganization model posits that neighborhoods with few economic resources will lack the political ability to resist sitings of toxic waste plants, the economic disadvantage perspective argues that those with the fewest economic resources will be *pushed into* the least desirable neighborhoods, which will tend to have the greatest proximity to toxic waste. Although some research has found that tracts with higher levels of income indeed have less proximity to such toxic sites (Bolin et al. 2002; Downey 1998; Hockman, and Morris 1998; Krieg 1995; Mohai, and Saha 2006), others have failed to find such a relationship (Baden, and Coursey 2002; Been 1995; Boer et al. 1997; Davidson, and Anderton 2000). As a consequence, a more nuanced viewpoint argues that the political power of high-income neighborhoods will allow them to resist toxic waste site placement, while the lowest income neighborhoods will also have few toxic waste sites due to their undesirable work force (Boer et al. 1997; Pastor et al. 2001; Sadd et al. 1999). This viewpoint suggests that working-class neighborhoods will have the greatest number of toxic waste sites.

Finally, while studies have focused on whether the economic resources of a neighborhood can help resist such sitings, prior research has rarely considered the possibility that residents' education level has important implications for weighing the risks of living close to a toxic waste site. The motivating insight here is that there is no reason to assume that perceptions of risk regarding toxic waste siting are both similar and high across population groups. Although most residents prefer not to live in a neighborhood with a plant emitting waste products, residents likely do not weight this risk equally. For instance, one study found that minority residents, specifically African Americans, were less likely than whites to move out of an area once a toxic waste plant had been sited there (Bullard et al. 2007). Hence, it is possible that there are differences in how those in various communities perceive the risks of living close to a hazardous waste site, and that the education level of residents may be a factor amplifying or attenuating risk perceptions of living close to waste sites.

Although it is important to understand the placement of toxic sites and the consequent exposure to toxic waste in a longitudinal framework, most studies utilize cross-sectional data. There are only a few exceptions to the general trend. For instance, by employing a case study of Los Angeles, Pulido (2000) argued that systematic racism operating through residential mobility patterns could explain disproportionate toxic exposure. Likewise, Bolin, Grineski, and Collins (2005) described a historical process that gave rise to segregation and environmental injustice in Phoenix, AZ. Such historical case studies provide a theoretical framework suggesting how such change might occur over time, but are limited by their single cases to making more generalizable claims. Likewise, a study providing descriptive information on how census tracts in Houston changed over time moves beyond a static perspective, but the lack of statistical modeling prevents firmer conclusions (Liu 1997). One study with longitudinal data on toxic sites did not actually estimate longitudinal statistical models, but instead focused on individual cases over time (Been 1994). One rare study attempting to tease out the causal direction between the placement of toxic sites and the presence of racial/ethnic minorities focused on tracts within the city of Los Angeles over three decades (Pastor et al. 2001). This limited literature suggests an important need for further longitudinal studies.

Our study addresses the following questions: 1) what is the relative proximity to toxic waste sites for minority groups over a ten year period from 1990 to 2000; 2) do neighborhoods with more highly educated residents experience less proximity to toxic waste sites; and 3) do these effects differ if we take into account the toxicity of the emitted wastes (Ash, and Fetter 2004; Brooks, and Sethi 1997; Neumann, Forman, and Rothlein 1998; Sicotte, and Swanson 2007). Beyond our longitudinal approach, we assess the impact of toxic sites on neighborhood residents by measuring the pounds of release weighted by a measure of its toxicity, and apportioning this value to a one-mile circle around the site rather than simply attributing it to the census tract in which the site is located (other studies using this approach include Bolin et al. 2002; Mohai, and Saha 2006). This third question is particularly important, as the actual toxicity of the waste is likely more highly correlated with the degree of health risk it poses. The present study focuses on six very highly populous counties in southern California. These six counties encompassed over 17 million persons in 1990, providing considerable insight into the process of proximity to toxic waste in our study sub-populations. The high proportion of Latinos in the study sample and relative proximity of this sample to the U.S./Mexico border provides insight into the experiences faced by immigrants to the U.S. Thus, our study population allows viewing the relative proximity to toxic waste for the burgeoning Latino population in this region over a ten-year period.

Data and Measures

Data for the present study were obtained from the U.S. Census and the Toxic Release Inventory (TRI), which contains information on the quantity of certain chemicals released into the environment by toxic waste facilities in the U.S. (obtained at http://www.rtknet.org/triabout.html). While the limitations of the TRI database are well-known--i.e., it only captures emitters of large amounts of toxic chemicals as facilities are required to file with the EPA each year if they exceed a threshold in release particulates of 25,000 pounds per calendar year for manufacture or processing activities, and 10,000 pounds per calendar year for “otherwise used” activities--it does give a reasonable approximation of the amount of such activity in a neighborhood (United States Environmental Protection Agency 2005). The U.S. Environmental Protection Agency (EPA) maintains this database. Although we are only focusing on one type of toxic waste emitter, some prior work suggests that differential proximity to various types of waste is similar (Bolin et al. 2002).[1]

Dependent variables

Our key outcome measures are the pounds of toxic waste emitted in a census tract per year over a decade, weighted by the inhalation toxicity. A census tract is a geographic area with a relatively homogeneous population of approximately 4,000 persons. The procedure for creating these measures is as follows. First, for each site, we geocoded its location. Second, we drew one-mile buffers around each site. While it is well-known that the area impacted by such toxic waste is variable based on the chemical involved and the wind patterns in the local area, prior studies have employed one-mile buffers as a reasonable approximation of the geographic dispersion of its impact. Indeed, one study used quarter mile and one mile buffers and found similar results for the two cutoff values (Pastor et al. 2001), whereas another study tested half mile, one mile and 2.5 mile buffers and also found similar results (Pastor et al. 2004). We estimated ancillary models using half mile and 2.5 mile buffers, and found substantively similar results to those presented using the one mile buffers. Third, we apportion this toxic waste based on the proportion of this buffer within each tract. Adopting this approach is important, as evidence suggests that sites are frequently on the border of tracts and thus realistically affect the health of residents outside their own tract (Mohai, and Saha 2006; Pastor et al. 2001). Finally, we take into account the toxicity of the particular chemical being released by multiplying the pounds of the chemical released by an inhalation toxicity score constructed by the Risk Screening Environmental Indicators (RSEI) study conducted by the United States Environmental Protection Agency (2004).

Independent variables

We constructed several tract-level measures with 1990 Census data. We accounted for the racial/ethnic composition of the tract by including measures of the percent African-American, Latino, Asian, and other race (using the percent white as the reference category). Prior research argues that the political power of high-income neighborhoods will allow them to resist toxic waste site placement, while the lowest income neighborhoods will also have few toxic waste sites due to their undesirable work force; we therefore accounted for this possible nonlinear effect by including measures of the average income in the tract and average income squared.

We measured residential stability as the average length of residence in the tract among all residents. We measured ethnic churning (EC) –the degree to which a tract undergoes racial/ethnic change in the prior decade—in a tract *k* by:

$$EC_k = \sqrt{\sum_{1}^{J} (G_{jt} - G_{jt-1})^2} \qquad (1)$$

where G represents the proportion of the population of ethnic group *j* out of *J* ethnic groups at time *t* (1990) and time *t-1* (1980). This gives a measure of the degree of racial/ethnic transformation that occurred in the tract in the decade leading up our study period: this is a sum of squares of differences, and we take the square root to return it approximately to the original metric. Thus, if there is no change in the racial/ethnic composition, it will have a value of zero. Finally, we measured the education level as the percentage with at least a bachelor's degree in the tract to capture highly educated persons.

Given that this is a sample of census tracts located in physical space, we took into account spatial effects in ancillary models.[2] There are three possible forms of spatial effects: 1) a spatial autocorrelation (or, error) effect, 2) a spatial lag effect of the outcome measure; 3) a spatial lag of the predictor measures. Ignoring #1 will affect the standard errors, and generally inflates the standard errors, suggesting that our tests here would be somewhat conservative (Anselin 2002). Although researchers frequently specify #2, recent scholars have called into question the wisdom of always employing such a default specification without more careful theoretical consideration (Elffers 2003; Morenoff 2003). We therefore follow the suggestion of Elffers (2003) and Anselin (2003: 161), among others, in specifying a model in which we test whether the spatial-lagged versions of our structural measures also impact tract toxic waste exposure.

Estimating spatial effects requires specifying what constitutes "close" neighborhoods, and we adopted a distance decay function with a cutoff at two miles (beyond which the tracts have a value of zero in the W matrix) in measuring the distance of surrounding tracts from the focal tract.[3] This resulting weight matrix (W) was then row-standardized—that is, each element is divided by the row total. We created spatially lagged measures by multiplying the values of our predictor variables by this W matrix. We estimated models both with and without these spatial measures to assess their impact. We tested for and found no excess collinearity among our independent variables in the main models (though the ancillary models including the spatial lags unsurprisingly showed evidence of increased standard errors and variance inflation factors around 7). The summary statistics for the variables used in the analyses are shown in Table 1.

<<<Table 1 about here>>>

Methods

Since we are interested in modeling the change over time of toxic waste proximity for various tracts in these six southern California counties, a latent trajectory model is an ideal strategy. Latent trajectory models use the observed values of the outcome variable at each time point to create an estimate of the latent trajectory over the time period (for a complete introduction to these models, see Bollen, and Curran 2006). This approach can use various linear and nonlinear specifications over time, and allows modeling the variability in these trajectories across the units of observation. We are modeling the trajectory of toxic waste in each of our study tracts over the 1990-2000 time period, which implies the following equation:

(2) $$y_{tjj} = \alpha_{ij} + \lambda_t\beta_{ij} + \lambda_t^2\beta^2_{ij} + \varepsilon_{tij}$$

where y is the toxicity-weighted amount of toxic waste at time *t* in tract *i* in county *j*, α is a random intercept that varies over tracts within a county, β is a random slope that varies over tracts in a county and has a λ effect on *y* (where lambda is structured to take into account time and therefore not estimated), β^2 is an optional random quadratic term that varies over tracts in a county with a λ^2 effect on *y* (and thus the λ's are not estimated), and ε is a disturbance, or error, term for each tract at each time point (which is assumed to be normally distributed with a mean of 0). Note that these random terms (α, β, β^2) are identified given that they are latent variables with loadings (the λ's) that are not estimated, but instead set to fixed values to capture time (for a complete discussion of the identification of such models, see Bollen, and Curran 2006: 21-25). Although various nonlinear specifications are possible, we found that a quadratic trajectory considerably improved the fit of the model compared to a linear trajectory and that a cubic trajectory only showed a modest improvement. Since we are most interested in a relatively parsimonious function of the change in toxic proximity in these tracts over time we utilized the quadratic trajectory in our analyses. This quadratic model showed a satisfactory fit based on the usual model fit indicators, with an RMSEA = .049 (values less than .05 are generally considered satisfactory), and values of .901 for the Comparative Fit Indicator and .905 for the Tucker-Lewis Indicator (greater than .90 is generally considered satisfactory for these measures).

We are interested in explaining the statistically significant variability in these random intercept and slope terms. This implies second-level equations such that:

(3) $$\alpha_{ij} = \kappa_\alpha + \Gamma_\alpha X_{ij} + \zeta_{1ij}$$

(4) $$\beta_{ij} = \kappa_\beta + \Gamma_\beta X_{ij} + \zeta_{2ij}$$

(5) $$\beta^2_{ij} = \kappa_\beta^2 + \Gamma_\beta^2 X_{ij} + \zeta_{3ij}$$

where α, β and β^2 are as defined before, the κ's represent the fixed intercepts for these random terms, X is a matrix of our tract-level predictors of interest (described in the data section) which has Γ_α effect on the random intercept (the amount of toxic waste exposure in the tract in 1990), a Γ_β effect on the random slope and Γ_β^2 effect on the random quadratic term (capturing the change in toxic waste exposure in the tract from 1990-2000), and the ζ's are disturbance terms assumed to have normal distributions with means of 0. We constrained these parameter estimates equal over counties given that the model fit was not appreciably reduced when doing so (the Bayesian Information Criterion was 289,783 for the model with constrained coefficients versus 290,528 for the model with parameters estimated freely over counties, where smaller values indicated better fit). We point out that the fact that there is likely additional random noise in the y_{tij} values given the uncertainty in the TRI data does not affect our parameter estimates in equations 3, 4, and 5.[4]

Since there may be differences across counties in exposure to toxic waste, we took these possible county processes into account by simultaneously estimating separate equations for each county (a multiple groups analysis). Thus, we are essentially viewing the effect of the tract-level variables on the proximity to toxic waste *compared to other tracts in the same county*. We account for the spatial differences among minorities with our tract and spatially weighted measures described above.

Results

To assess the toxic proximity for residents in the tracts of these counties, we begin by examining the models that do not include our tract-level predictors. Although the model estimates a trajectory for each tract in the sample, we visually plot the *average* trajectory for the tracts in each county in Figure 1. Note that all of these linear and quadratic slope terms were significantly different than zero, except those for San Bernardino. The level of toxic waste proximity generally declined over this ten-year period. However, there are some differences across counties: whereas tracts in Los Angeles and Orange counties generally showed declining exposure to toxic waste over this period, and the tracts in Ventura County showed a general decline with a slight uptick towards the end of the decade, the tracts in the other counties followed different trajectories of exposure. Whereas tracts in San Bernardino County showed little trend over this period, tracts in San Diego and Riverside counties showed a modest decrease in exposure to toxic waste over the decade.

<<<Figure 1 about here>>>

We next attempt to explain *why* some tracts have higher levels of *relative* toxic waste proximity in 1990 (the random intercept), and why some have differing trajectories. These are the coefficients from equations 3, 4, and 5. In these models, the first column of Table 2 highlights that, holding all other measures in the model constant, tracts with a higher percentage of Latinos or Asians have higher levels of toxic proximity in 1990 (compared to tracts with higher levels of white residents), though tracts with African Americans have *lower* exposure than whites. We also see that tracts with higher levels of residential instability and ethnic churning during the 1980s have higher levels of toxic proximity in 1990. The quadratic effect for average income indicates a curvilinear effect in which tracts with the highest and lowest levels of levels of income have lower levels of toxic proximity, whereas tracts with middle class residents experience the greatest toxic proximity. Finally, tracts with more highly educated residents have significantly less toxic waste in 1990.

<<<Table 2 about here>>>

The model in Table 2 also shows that these racial/ethnic groups experience different trajectories in toxic waste proximity over this decade. For example, whereas tracts with a higher percentage African-American at the beginning of the decade have lower toxic exposure such tracts experience a greater increase in toxic proximity over this decade (as seen by the significantly positive effect in column 2 predicting the random linear term). This increased proximity slows over the decade (as seen by the significant negative term for percent African-Americans in the third column predicting the random quadratic term). The other findings in this model can be interpreted similarly: for instance, tracts with higher levels of racial/ethnic churning in the prior decade experienced a greater decrease in toxic waste proximity over this decade, as did tracts with a more highly educated population.

To ascertain the magnitude of these effects, we input the estimates from Table 2 into equations 2, 3, 4, and 5 to get estimated values at each time point for these various subpopulations. That is, we first computed the level of toxic waste proximity in a tract with average values for all of our predictor variables, then computed the value for a hypothetical tract in which the variable of interest was increased one standard deviation, and then computed the percentage change. For the racial/ethnic variables, we calculated this as a 15 percentage point increase for each to keep them in a common metric (since their differing standard deviations are a function of this particular sample). This provides a sense of the degree of differential proximity to this toxic waste that exists in the study. We plotted the percentage increase or decrease of these values compared to an average tract in Figure 2.

<<<Figure 2 about here>>>

We see in Figure 2 that in 1990, a 15 percentage point increase in African-Americans in a tract that is average on all other characteristics results in 33% less toxic waste proximity in 1990. This same tract, however, experienced a steady increase in toxic waste exposure over the immediately subsequent years, such that these residents had 12.6% more toxic waste proximity by 1994, 17.1% more proximity in 1995, up to 18.9% more proximity in 1996. From that point, the negative effect of percentage African-Americans on the random quadratic term shows its effect as the gap between this tract and an average tract narrows: by 2000, this tract with 15 percentage points more African-Americans experiences an average proximity to toxic waste. Thus, on average, this tract with more African-Americans will experience 3.2% more proximity to these highly toxic wastes over the study period than will an average tract.

The effects are much stronger for Latinos and Asians, as a 15 percentage point increase in these groups in a tract with average characteristics results in a 59 to 78% increase in proximity to toxic waste in 1990. But whereas a tract with more Asians shows a steady improvement over the decade—such a tract has only 7.5% more exposure to such toxic waste by 2000—a tract with more Latinos retains its relatively high exposure over this decade. Thus, a tract with 15 percentage points more Latinos is exposed to, on average, about 84% more toxic waste than an average tract over this period. And whereas a tract with more Asians is exposed to about 34% more toxic waste over this period, at least this differential has nearly dissipated by the end of the decade.

Turning to the two social disorganization measures, we see that residential instability and ethnic churning both increase the proximity to toxic waste. A one standard deviation increase in ethnic churning in the previous decade leads to a 28% increase in proximity to toxic waste in 1990. While this difference lessens in the early part of the decade, by the latter half of the decade conditions are similar to those at the beginning of the decade. And tracts with one standard deviation more instability go from about 34% more toxic exposure in 1990 to almost 50% more by 2000.

There are also differential effects for the SES measures. The curvilinear effect for average income complicates interpretation, so we do not plot it here. However, there are strong effects for the presence of highly educated residents. Given the concern that the effect of educated residents may overlap with the income effect (as their effects were of opposite sign), we plot here the education effects from a model that did not account the average income of the residents. Although the size of the effect here is about 30% smaller than the model including average household income, there are still quite dramatic effects as a one standard deviation increase in the percentage with a bachelor's degree (about 15.5%) in an average tract decreases proximity to toxic waste 62% in 1990. This advantage improves further over much of the decade: by 1997 such a tract is exposed to almost 100% less toxic waste. Thus, over the years of this decade, this increase in the highly educated residents in a tract results in about 89% less proximity to toxic waste than an average tract.

Raw pounds of output

We next tested the importance of taking into account the toxicity of the wastes being emitted. We accomplished this by estimating models in which we substituted the raw pounds of toxic wastes emitted for our toxicity-weighted measure. While it is not reasonable to compare the size of unstandardized coefficients from the models given the different metrics, we can compare the relative size of the *standardized* coefficients. We plot the relative comparison of the standardized coefficients from our models accounting for the toxicity of the emitted waste to the model only accounting for the raw pounds of waste in Figure 3. We do not plot the differences for coefficients that were not significant in either of the models.

<<<Figure 3 about here>>>

It is notable that some of the findings are considerably stronger when taking into account the toxicity of the wastes. For example, the left side of Figure 3 shows that the size of the standardized effect of percent African Americans on the random slope is 34% stronger in the model accounting for the toxicity than the model predicting just raw pounds, and the effect on the random quadratic term is 67% stronger (though there is no difference on the random intercept). For Latinos, the effect on the random intercept is 65% stronger when we account for the toxicity of the releases. Likewise, the effects on the random intercept are 57% stronger for residential stability and about 40% stronger for the average household income measures when accounting for the toxicity. We also see an important difference for Asians, as the effect on the random slope is only 30% as large in the model accounting for toxicity: thus, whereas the situation in the neighborhoods Asians live in is improving considerably when measuring this based on the toxicity of the waste, this effect would not be detected if we only used the raw pounds of waste. The differences are also pronounced for racial/ethnic churning: whereas the size of the standardized effect on the random intercept is about 30% smaller when accounting for the toxicity levels, it is actually 34 and 100% larger on the random slope and quadratic terms. Finally, the advantaged position of neighborhoods with more highly educated residents is accentuated when accounting for the toxicity of the waste: the size of the standardized coefficients are 18% stronger for the random intercept and about 33% stronger for the random slope and quadratic terms.

Spatial effects

Finally, we briefly discuss ancillary models including the spatially lagged terms. These spatially lagged measures introduced a considerable degree of collinearity, and thus imprecision, into the estimates. Nonetheless, the pattern of results was little changed (see Table A1 in the Appendix). Among the spatially lagged variables, the only significant effects for the model predicting the random intercept at the beginning of the time period were the spatial lags for Latinos and educated residents. Specifically, the presence of highly educated residents in nearby tracts is even more important for the presence of toxic waste in a tract than the presence of highly educated residents in the tract itself. This might suggest that such residents are not just concerned about toxic waste within their own tract, but also placement in nearby tracts as well. And although there is more toxic waste in tracts with more Latinos, such tracts have less toxic waste when there are more Latinos residing in nearby tracts. It therefore appears that tracts located in the center of a barrio are somewhat cushioned from the presence of such toxic waste.

In the equations predicting change over time, the only significant effect was that the presence of high income households in nearby tracts soaks up some of the effect we observed of high income households in the tract itself. That is, whereas tracts with more high income residents actually experienced an increase in toxicity over time, this effect was actually enhanced if they were located in areas surrounded by other high income tracts. (There was no evidence of a curvilinear effect for average tract income or average income in the surrounding tracts in this model).

Discussion

Although prior research has frequently cross-sectionally investigated the effect of disproportionate proximity for different populations to toxic waste, our study explored how this proximity can change over time. Our findings provide several important insights. Our results showed that the burgeoning Latino population in the region wherein the study population resides experienced dramatically more relative proximity to toxic waste compared to other minority groups. Hence, our study offers insights into the experiences impacting the health of Mexican immigrants to the U.S. Second, by employing longitudinal data over a ten-year period, we moved beyond prior work viewing the co-occurrence of various sub-populations and toxic waste sites and estimated the relative proximity to toxicity-weighted waste for various groups, and how this changed over this ten-year period from 1990-2000. Third, we measured the toxic waste in the communities under study by accounting for the actual location of a toxic site (and a one mile buffer around it) rather than simply attributing all of its output to the census tract in which it is actually located. Moreover, we also accounted for both the *amount* of toxic waste that is emitted (Perlin et al. 1995) and its *toxicity*, rather than simply measuring the presence of such a site. A key finding of the present study is that tracts with many highly educated residents have particularly low, and declining, proximity to such toxic waste. Although we found a general decline in proximity to toxic waste over this decade, subpopulations are disproportionately exposed and experienced differential declines over the decade.

Our findings emphasize the importance of taking into account the toxicity of the waste being emitted by these plants (Ash, and Fetter 2004; Brooks, and Sethi 1997; Neumann et al. 1998; Sicotte, and Swanson 2007). Whereas prior work frequently views only the presence of a toxic waste site in a tract, or the raw number of pounds of waste being emitted, differential effects were found when taking into account the relative toxicity of the emitted waste. We often found that the effects are much stronger (based on standardized coefficients) when accounting for the toxicity of the waste. For example, whereas tracts with more Latinos experienced more raw pounds of waste effluence than did tracts with more whites, the size of this difference is greater when comparing proximity based on the toxicity of the chemicals. Similarly, neighborhoods with more highly educated residents not only experience less toxic waste as measured based on raw pounds, but they are particularly likely to experience lower levels of the most toxic waste based on the RSEI toxicity scale. Of course, not all toxic substances are equally toxic, and hence equally harmful, to all persons. Although the toxicity of various compounds can vary depending on the particular subpopulation that is exposed to it (depending on such characteristics as the age and body mass of the person), it is nonetheless important to take such toxicity differences into account as studies have suggested that cumulative exposure is particularly harmful (Faber, and Krieg 2002; Mott 1995; United States Environmental Protection Agency 2004).

There was evidence consistent with social disorganization hypothesis that tracts undergoing racial/ethnic transformation or general instability are most likely to experience greater proximity to such highly toxic wastes. This might suggest that these tracts are politically disenfranchised and possibly fragmented in terms of political power given that they likely do not have stable cores of residents that might act collectively to initiate local policy movements to prohibit toxic waste siting or to insist on reductions in the level of toxic waste emitted from already existing plants. On the other hand, tracts with the lowest levels of income experienced *lower* levels of proximity to toxic waste over this time period, in contrast to the social disorganization hypothesis. This may suggest that high income tracts have the political ability to resist such sitings, and that working-class neighborhoods are more likely locations. Nonetheless, future research will need to test whether these mechanisms indeed explain the pattern of results found here.

Our findings emphasized the importance of accounting for the education level of the residents in understanding how disproportionate proximity affects populations. Tracts with a higher proportion of highly educated residents were exposed to lower levels of toxic waste in 1990, and this proximity declined even further over the study period. These findings suggest that it is the education of residents—rather than their economic resources—that explain which neighborhoods will be most impacted by the placement of toxic waste sites. This is consistent with prior work suggesting that more educated residents have lower risk thresholds (Gaba, and Viscusi 1998). These findings for educated residents were even stronger when taking into account the level of toxicity of the chemicals being emitted. This may occur because these residents are more aware of the health risks represented by highly toxic wastes and therefore avoid such neighborhoods, or because they are more willing to engage in activity to prevent the placement of such sites and to reduce the output from currently existing sites, particularly when it comes to the most toxic chemicals. This is clearly speculative, suggesting a useful direction for future research.

Likewise, our finding of disproportionate proximity to toxic waste among Latinos may suggest a limited awareness among this largely immigrant population about the risks involved in living near such toxic waste sites. It appears that the highest toxicity chemicals—those that are arguably the most harmful—are precisely the ones to which Latinos remain disproportionately exposed. Given the evidence that Latinos are no less likely to perceive risks of which *they are aware* (Williams, and Florez 2002), and that they engage in self-protective behavior when they are aware of risks (Vaughan 1993), our findings may suggest a lack of awareness and knowledge among Latinos pertaining to the actual toxicity of wastes, extent of proximity, and possible deleterious consequences on health that could follow.

Limitations

Although this study has provided considerable insight into the disproportionate proximity of some subpopulations to toxic waste, some limitations should be noted. First, although we took into account the level of waste emitted by these plants and its toxicity, it is nonetheless the case that these are approximations that are affected by actual wind patterns in areas as well as the individual characteristics of residents. Nonetheless, although one study looked at the effect of wind patterns on the waste of one plant in Oakland (Fisher, Kelly, and Romm 2006), we know of no empirical research showing that wind patterns systematically favor the communities in which minorities reside. Second, it is well known that our data source does not contain all toxic waste emitters. Given that these are particularly large emitters, we suggest that these patterns are relatively representative of the sort of cumulative exposure experienced by these subpopulations.

Nonetheless, there are other sources of pollution to which populations are exposed, and we make no claim that we are measuring all such exposures. Third, although we have extended the literature by exploring a longitudinal model, it is nonetheless the case that we are limited in our ability to provide causal explanations. Fourth, we were limited by data availability to focus on one specific ten-year period. Clearly, future research will need to study these patterns over longer periods to gain further insights. Finally, our measure of highly educated residents is of necessity a somewhat crude measure. We would certainly have preferred to have information on the level of education and knowledge specific to toxic waste and its consequences.

Implications

Study findings suggest the merit of educating minority populations about the health risks associated with toxic output and living in close proximity to toxic sites. Efforts including educational outreach programs and community wide campaigns tailored to raise awareness among minority populations may have the potential to increase knowledge concerning the risks of living in close proximity to hazardous waste sites and may be instrumental in lessening the migration of such populations into neighborhoods with hazardous sites. Educating those in vulnerable communities of the health risks of living proximally to such plants may ignite community organization strategies for social action to combat injustices or actively resist the placement of such sites. Such grassroots initiatives can also lead residents to band together to minimize the level of output of existing sites, which could also limit the disproportionate proximity of certain groups.

If grassroots initiatives were to create a sentiment of reduced willingness of residents to move to such neighborhoods, housing costs may start to fall in these neighborhoods. This reduced housing cost might be perceived as an additional externality caused by the emitting plant, and lead to increased calls for restitution to the neighborhood. This would suggest that such sites are not currently contributing their fair share given the negative externalities to which they are exposing such neighborhoods. To the extent that residents are made more aware of the health risks involved in such exposure, the health costs imposed by such plants on communities will be more accurately judged, more equally borne by subgroups, and more equitably dealt with by urban planning commissions. One overall consequence might be a general focus on limiting emissions.

Efforts at the community level may be the impetus for larger scale policy changes at the federal level that may more strictly ensure the proper identification, assessment, evaluation, and regulation of toxic waste exposure among populations in vulnerable communities. Patterning on the work of existing community and federal partnerships, such as the Mississippi Delta Project, a model or interorganizational collaboration in the area of environmental justice (Agency for Toxic Substances and Disease Registry 1998) should be looked to for guidance on how to create synergistic and effective interorganizational networks of local, state, and federal partners to use a public health systems approach to combat environmental injustices and improve the health of minority populations in southern California.

Conclusion

Disproportionate proximity to toxic waste is a significant public health issue that likely contributes to the disproportionate disease burden among minority populations in southern California. A fundamental message from this research is a possible role that education provided by health professionals could play in alleviating differential proximity. The evolution of grassroots efforts targeting disadvantaged communities that eventually lead to local, state and federal level policies that enact stricter regulatory laws relating to the siting of toxic waste are critical for combating the environmental injustices differentially burdening the health of minority populations in southern California today.

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Endnotes

[1] In prior years the quality of the addresses were admittedly relatively poor (sometimes including addresses of head offices instead of the actual facilities), but work in recent years has improved the quality of the addresses in these data (United States Environmental Protection Agency 2007).

[2] Furthermore, ignoring a spatial lag effect of the outcome measure if this is in fact the true relationship (in which the toxic waste in one neighborhood increases toxic waste in adjacent neighborhoods), would result in our estimation of the *total* effects of our predictors, rather than *direct* effects (Anselin 2003). That is, the presence of minorities in a tract may be associated with higher levels of toxic waste, and these higher levels of toxic waste then impact the amount of toxic waste in adjacent tracts. This implies that the presence of minorities in one tract *indirectly* increase the amount of toxic waste in adjacent tracts. Since our goal is not to parse apart these direct and indirect effects, we suggest that these total effects are of most interest in understanding the effect of toxic waste proximity for minority populations.

[3] We also estimated ancillary models using a one-mile distance decay function for the spatial weighting matrix, and a three-mile distance decay function. The substantive results were very similar (results available upon request).

[4] The reason for this may not be immediately apparent. In the case of a linear regression model, measurement error in the outcome variable does not bias the coefficients of the variables predicting that outcome, but only increases their standard errors (Greene 2000: 375-380). In our case, this affects equation 2. However, this additional uncertainty in equation 2 has no effect on the estimates predicting the latent variables in equations 3, 4, and 5. This is because the total error in equation 2 is parsed into uncertainty and regression error, but the total error remains the same. Indeed, when we estimated ancillary models in which we accounted for lower reliability of these measures (we set them at .70 reliability), our parameters estimates of interest and

standard errors in equations 3, 4, and 5 remained essentially unchanged. Setting the reliabilities to different values yields the exact same results.

Tables and Figures

Table 1. Summary statistics of toxicity-weighted effluent and socio-demographic variables in six counties in southern California. N = 3,001 census tracts

	Mean	Std. Dev
Weighted toxic waste, 1990	8.77	5.06
Weighted toxic waste, 1991	8.16	5.07
Weighted toxic waste, 1992	7.89	4.99
Weighted toxic waste, 1993	7.48	5.15
Weighted toxic waste, 1994	7.30	5.13
Weighted toxic waste, 1995	6.74	5.15
Weighted toxic waste, 1996	6.55	5.21
Weighted toxic waste, 1997	6.04	5.27
Weighted toxic waste, 1998	5.83	5.16
Weighted toxic waste, 1999	5.91	5.13
Weighted toxic waste, 2000	5.82	5.07
Percent African-American	7.93	15.54
Percent Latino	28.82	24.80
Percent Asian	8.67	9.66
Percent other race	0.66	0.78
Percent white	53.92	29.88
Ethnic churning	17.21	12.20
Residential stability (average length of residence)	9.04	3.13
Average family income ($1,000's)	52.81	29.35
Percent with at least a bachelor's degree	12.62	10.94

Table 2. Trajectory of toxic wastes (pounds of waste weighted by inhalation toxicity, logged) for census tracts in six counties in southern California. N = 3,001 census tracts

	Random intercept			Random slope			Random quadratic		
	Coef.	SE	Sig	Coef.	SE	Sig	Coef.	SE	Sig
% African-American	-0.017	(0.008)	*	0.087	(0.024)	**	-0.071	(0.020)	**
% Latino	0.029	(0.007)	**	0.012	(0.019)		-0.011	(0.016)	
% Asian	0.039	(0.012)	**	-0.031	(0.036)		-0.003	(0.030)	
% other race	-0.439	(0.152)	**	0.599	(0.381)		-0.243	(0.309)	
Racial/ethnic churning	0.022	(0.010)	*	-0.067	(0.030)	*	0.067	(0.025)	**
Residential stability	-0.108	(0.039)	**	0.028	(0.113)		-0.078	(0.096)	
Average family income	0.418	(0.107)	**	0.501	(0.243)	*	-0.400	(0.206)	†
Average family income squared	-0.043	(0.011)	**						
% with a bachelor's degree	-0.059	(0.014)	**	-0.113	(0.040)	**	0.083	(0.033)	*

*** p < .01(two-tail test), * p < .05 (two-tail test), † p < .05 (one-tail test). Standard errors in parentheses.*

Figure 1. Unconditional trajectories of toxic waste (weighted by toxicity) in six southern California counties. N = 3,001 census tracts

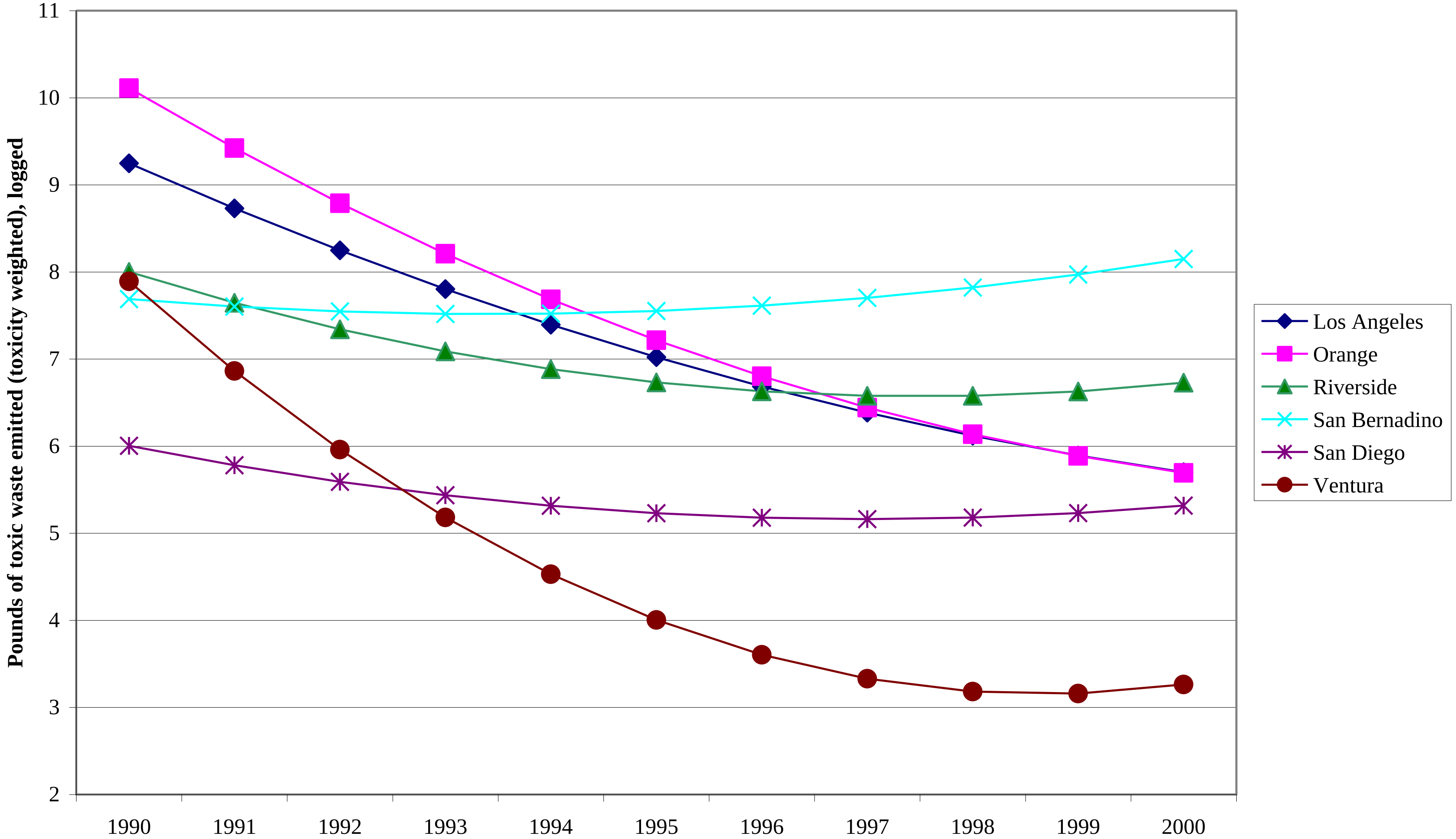


Figure 2. Percentage difference in exposure to toxicity-weighted wastes between average tracts in sample and tracts one standard deviation higher on selected variable (15% higher for racial/ethnic groups). N = 3,001 census tracts

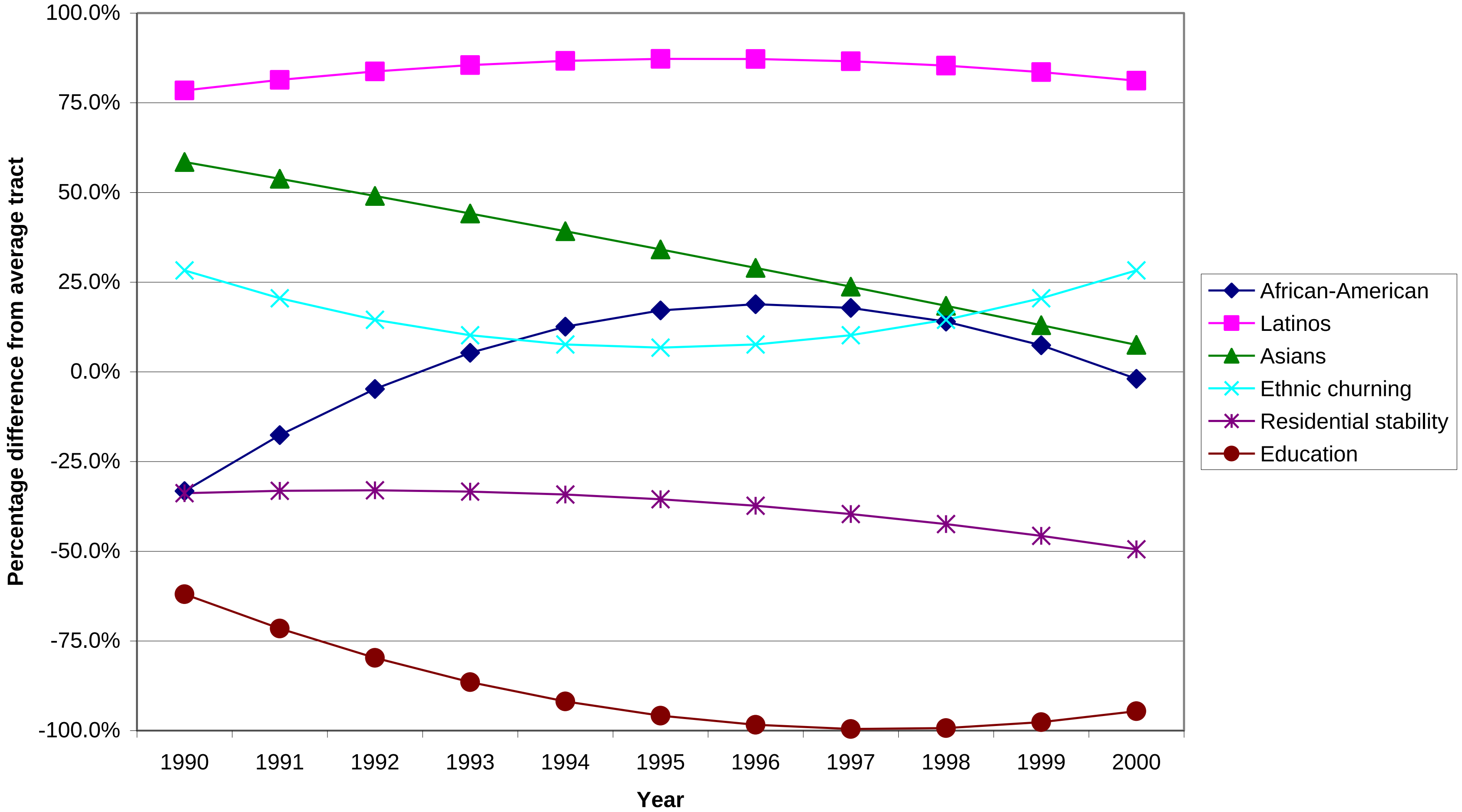
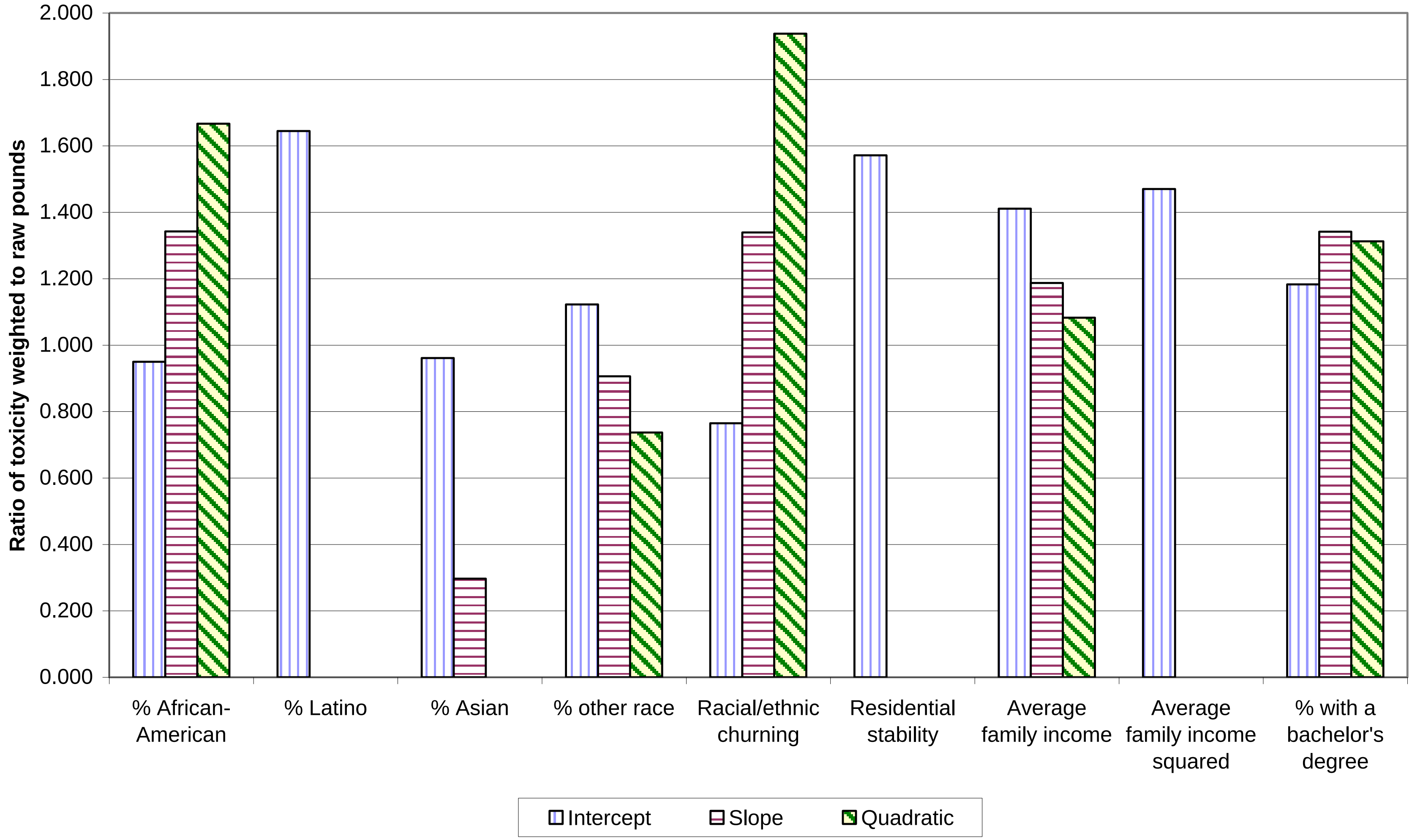


Figure 3. Ratio of the standardized coefficient estimates from the model with toxicity-weighted measure to the standardized coefficient estimates from the model with raw pounds measure



Appendix

Table A1. Trajectory of toxic wastes (pounds of waste weighted by inhalation toxicity, logged) for census tracts in six counties in southern California. Accounting for spatial lag effects of predictor variables. N = 3,001 census tracts

	Random intercept			Random slope			Random quadratic		
	Coef.	SE	Sig	Coef.	SE	Sig	Coef.	SE	Sig
% African-American	-0.020	(0.017)		0.119	(0.052)	*	-0.094	(0.044)	*
% Latino	0.020	(0.010)	†	0.030	(0.029)		-0.040	(0.024)	
% Asian	0.022	(0.012)	†	-0.006	(0.037)		-0.028	(0.031)	
% other race	-0.564	(0.166)	**	0.738	(0.384)	†	-0.386	(0.314)	
Racial/ethnic churning	0.007	(0.010)		-0.067	(0.030)	*	0.066	(0.026)	*
Residential stability	-0.095	(0.046)	*	0.053	(0.129)		-0.068	(0.108)	
Average family income	0.086	(0.099)		0.151	(0.282)		-0.174	(0.237)	
% with a bachelor's degree	0.016	(0.019)		-0.100	(0.052)	†	0.077	(0.044)	†
Spatial lags									
% African-American	-0.026	(0.020)		-0.020	(0.061)		0.013	(0.051)	
% Latino	-0.030	(0.013)	*	-0.005	(0.037)		0.022	(0.031)	
Residential stability	-0.032	(0.076)		0.106	(0.215)		-0.233	(0.179)	
Average family income	-0.106	(0.174)		1.226	(0.489)	*	-0.902	(0.408)	*
% with a bachelor's degree	-0.135	(0.028)	**	-0.068	(0.077)		0.044	(0.063)	

*** p < .01(two-tail test), * p < .05 (two-tail test), † p < .05 (one-tail test). Standard errors in parentheses.*