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The Pen and the Scalpel: Effect of Diffusion of Information on Non-Clinical Variations in Surgical Treatment

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Abstract

Background—As information is disseminated about best practices, variations in patterns of care should diminish over time.

Objective—To test the hypotheses that differences in rates of a surgical procedure are associated with type of insurance in an era of evolving practice guidelines and that insurance and site differences diminish with time as consensus guidelines disseminate among the medical community.

Methods—We use lymph node dissection among women with ductal carcinoma in situ (DCIS) as an example of a procedure with uncertain benefit. Using a sample of 1051 women diagnosed from 1985 through 2000 at two geographic sites, we collected detailed demographic, clinical, pathologic, and treatment information through abstraction of multiple medical records. We specified multivariate logistic models with flexible functions of time and time interactions with insurance and treatment site to test hypotheses.

Results—Lymph node dissection rates varied significantly according to site of treatment and insurance status after controlling for clinical, pathological, treatment, and demographic characteristics. Rates of lymph node dissection decreased over time, and differences in lymph node dissection rates according to site and generosity of insurance were no longer significant by the end of the study period.

Conclusions—We have demonstrated that rates of a discretionary surgical procedure differ according to non-clinical factors, such as treatment site and type of insurance, and that such unwarranted variation decreases over time with diminishing uncertainty and in an era of diffusion of clinical guidelines.

Keywords

diffusion; insurance; variation; ductal carcinoma in situ

Introduction

Explanations for variation in surgical procedures include medical uncertainty about the merit of a particular treatment,^{1–3} inappropriate use in regions with high rates of procedures (supplier-induced demand),¹ random variation,² enthusiasm for particular services,⁴ and defensive medicine.^{2, 5} Diffusion theory suggests that uncertainty regarding the merits of a treatment leads to variation in the use of that treatment.^{2, 6} As more knowledge is generated and disseminated about the value and negative consequences of a treatment, practice variations diminish,⁷ and practice patterns may ultimately converge. In addition, clinical uncertainty about the value of a treatment provides a window of opportunity for economic incentives, particularly reimbursement, to drive physician behavior.⁵ Even without clinical uncertainty, if treatment strategies have equivalent survival, the choice of treatment may be particularly sensitive to relative payments.^{8, 9}

The purpose of this study was to investigate the impact of evolving practice guidelines on surgical management of the axilla in women with ductal carcinoma in situ (DCIS) and to investigate the role of physician financial incentives in practice patterns. Beginning in 1987, observational studies and expert opinion¹⁰ followed by textbooks,¹¹ and then published guidelines¹² advised against lymph node dissection in women with DCIS based on lack of benefit.

We hypothesized that lymph node dissection rates have been affected by the generosity of insurance during an era of guideline evolution and dissemination. If financial incentives play a role in use of a medical treatment, higher rates of the treatment would be expected among patients for whom the reimbursements are more favorable. We further hypothesized that differences in rates of lymph node dissection among women with different types of insurance would diminish with time as consensus about the value of a procedure was reached and disseminated among the medical community.

Methods

We conducted a retrospective study to examine patterns of care and outcomes in women with DCIS diagnosed between 1985 and 2000. The institutional review boards at the

University of Rochester, the Henry Ford Health System, and the RAND Corporation approved the study.

Conceptual Model—Our conceptual model starts with the Behavioral Model of Health Services Use developed by Andersen and Aday,^{13, 14} which incorporates patient, physician, and health system factors into an understanding of health care utilization. We thereby account for factors associated with the *clinical need* for lymph node dissection and other factors that facilitate or impede access to care (referred to as *enabling* and *predisposing* factors in the typology of Andersen).

The clinical need for lymph node assessment in women with DCIS is restricted to women with a high likelihood of invasive disease.^{10, 15, 16} Factors associated with a higher likelihood of invasive disease include a palpable mass,¹⁷ extensive disease on mammogram,¹⁵ multifocal disease,¹⁸ younger age,¹⁵ comedo histology,^{18, 19} and high histologic grade.¹⁵ In addition, in patients with a high likelihood of invasive disease who are undergoing mastectomy and autogenous reconstruction, lymph node assessment is particularly appropriate because lymph node dissection following autogenous breast reconstruction may compromise the blood supply to the reconstructed breast. The above disease, patient, and surgical treatment factors were thus included in our model as need characteristics.

Factors that can facilitate or impede access to care, such as race, income, insurance status, comorbidity, and site of care, have been shown to be associated with receipt of lymph node dissection in women with invasive disease^{20–25} (in whom lymph node assessment is required for staging and selection of locoregional and systemic treatment). We thus include these enabling and predisposing factors in our model because they may similarly play a role in receipt of lymph node dissection in women with DCIS.

Finally, we expand this framework to include our key hypothesis that diminishing uncertainty over time is associated with diminishing differences in practice by including year of treatment and year of treatment interacted with type of insurance and treatment site in the analytic model.

Study Sample—Study subjects were women diagnosed with DCIS between 1985 and 2000 within two tumor registries – the population-based Monroe County (New York) tumor registry and the tumor registry of the Henry Ford Health System. Exclusion criteria included invasive breast cancer (including microinvasive disease) or a history of a previous breast cancer. The final analytic sample contained 1051 patients.

Monroe County Sample: The Monroe County (New York) Tumor Registry was used to identify women diagnosed with DCIS during our study time frame who received care from 24 surgeons who had 10 or more patients in the Registry. The Registry identified 934 observations. Three observations were duplicate subjects, leaving 931 unique subjects. We were unable to locate charts for 92 subjects. In chart review we confirmed that 153 of the 839 subjects (18.2%) were not eligible because of a previous cancer history or concomitant invasive disease (that is, the patient did not have pure DCIS), leaving 686 subjects. We dropped an additional 21 subjects because of incomplete data, leaving 665 subjects with complete chart abstractions. In analyses following chart abstraction, we identified another 27 women (4.1% of women with complete data) who were not eligible because they had Paget disease or lobular carcinoma in situ, leaving 638 in our final analytic sample. Assuming a similar rate of ineligibility among the subjects not located, we estimate that there were 730 eligible women in our sampling frame. Our final analytic sample consists of 87.4% of the women we estimated to be eligible for inclusion in the study.

Henry Ford Health System Sample: The Henry Ford Health System electronic registry was used to identify women diagnosed with DCIS during our study time period. The registry identified 750 observations, 435 of whom were eligible for inclusion. We were able to local complete data for 421 (96.7%) patients. In analyses following chart abstraction, we dropped an additional 8 patients who were found to be ineligible (because of Paget disease or only lobular carcinoma in situ). As above, we assume a similar ineligibility rate among the women with incomplete records, resulting in a final analytic sample 96.8% of the eligible subjects.

Data Collection—The primary data collection instrument was completed by trained medical record abstractors who obtained data from an exhaustive review of surgical records, hospital records, radiation oncology records, medical oncology records, pathology reports, operative reports, and, when necessary, primary care and gynecology records. The data collection instrument included date of diagnosis, socioeconomic and demographic characteristics (date of birth, self-assigned race and ethnicity, and address for census block group assignment), disease and treatment characteristics (grade, size, width of the surgical margins, presence of a palpable mass at diagnosis, means of detection, mammographic extent of disease, type of surgery, number and type of surgical procedures, and type of lymph node assessment, if any), comorbid conditions included the Charlson comorbidity index,²⁶ menopausal status, and treating physician. Census level socioeconomic status was measured as percentage of people living below poverty and percentage of adults with at least a high school diploma at the census block group level from the 2000 Census. Operative and pathology reports were used to confirm that axillary lymph nodes were not removed unintentionally in patients having mastectomy. Continuous quality checks of the data were performed by a medical oncologist, surgical oncologist, and breast pathologist.

Statistical Analysis—We began by generating univariate and bivariate descriptive statistics. The bivariate analyses assess the relationship between a dichotomous indicator for lymph node dissection and each of the independent measures, including socioeconomic and demographic measures (year of diagnosis, age at diagnosis, race, insurance status), comorbidity, surgical treatment (breast conserving surgery or mastectomy and type of reconstruction), and disease characteristics (mammographic size, histologic subtype, and nuclear grade). We report Pearson's chi-squared tests of independence for categorical variables. We also report Person's chi-squared tests of independence for each dichotomous covariate defined for the multivariate model.

We estimated multivariate logistic regression models to describe the relationship between the independent measures and lymph node dissection. Following the discussion above, we included groups of independent variables to control for the characteristics thought to be related to *clinical need* (receipt of mastectomy, receipt of autogenous reconstruction, extensive or multifocal disease, age, presence of a palpable mass, comedo histology, and high grade histology), *predisposing/enabling* factors (race, income, insurance, site of care, and comorbidity), and characteristics related to our *key hypotheses* (calendar time and the interaction of calendar time with insurance and site of care). Medicare plus supplemental private insurance was assumed to be the most generous insurance and Medicaid the least. We included a site indicator and interactions of site and surgical treatment to identify differences in treatment patterns (rates of lymph node dissection) across our two sites by surgical treatment. Because our key hypotheses are related to patterns of lymph node dissection use over time, we specified flexible functions of time and interactions with time. We formulated a standard logistic regression model of the form

$$P(LND=1) = \frac{1}{1 + e^{\lambda\beta + \vartheta(t,Z)}}$$

where lymph node dissection is indicated as LND, X is a vector of covariates described above, Z includes insurance status and site indicators, t denotes calendar year, and the function $\vartheta(\cdot)$ flexibly models time dependence. We specify $\vartheta(\cdot)$ as overlapping polynomials, first introduced in the health literature by Garber and MaCurdy in 1993.²⁷ Unlike typical polynomial splines, overlapping polynomials allow functions in different intervals to overlap, turning on or off smoothly. We specified $\vartheta(\cdot)$ as

$$\vartheta(t, Z) = \sum_j \left[\alpha_1^j Z_j * \Phi\left(\frac{t - \mu_j}{\sigma_j}\right) + \alpha_2^j Z_j * \left(1 - \Phi\left(\frac{t - \mu_j}{\sigma_j}\right)\right) \right],$$

where j indexes the elements of Z , $\Phi(\cdot)$ is a standard normal cumulative distribution function, and the α 's are estimated coefficients. The location and smoothness of the overlapping polynomials are determined by μ (a mean shift) and σ (a standard deviation), respectively (see Appendix for illustrations). As σ approaches 0, the overlapping polynomial collapses to an indicator variable that turns on or off at the year given by μ . As σ increases, the overlapping polynomial turns on or off smoothly over time. For example, if we set $\mu = 1995$ and $\sigma = 2$, $\Phi((t-1995)/2)$ will be approximately zero for observations with calendar year < 1990 , it will gradually increase in value reaching 0.50 at 1995 and approximately 1 by 2000. In principle, we could estimate the values of the overlapping smooth-spline parameters (μ , σ). Here, however, we performed a grid search over pairs of μ , σ to maximize model fit. The resulting model estimates allow for insurance status and site to have smooth interactions with time, as determined by the data. The overlapping polynomials illustrated in the Appendix, as described above, show the characteristics of the functions with the optimal parameter values from our estimation.

Our socio-economic measures are based on census tract data. These include various measures of the distribution of race, language spoken at home, travel time to work, education, and income within census tract. Because these measures are highly correlated, we present specifications that include percent below poverty, and percent black in the census tract. In addition to these census tract-level variables, we also have individual-level data on race. By including both the census-based race measure and the individual-level race measure, we can identify the contribution of both neighborhood race effects and individual race effects.

We present results for a parsimonious model in which we drop covariates, one at a time, until all included measures have p-values < 0.20 . We do not, however, drop any of the measures that are related to our key hypotheses (treatment measures and insurance status measures). Results from this parsimonious model were not substantively different from results from a model that included all covariates.

After exploratory analyses for best functional form for calendar time, our final models specify calendar time as a second order polynomial (linear and quadratic terms). All analyses are performed using Stata version 9.0 (StataCorp LP, College Station, TX).

Results

Table 1 presents descriptive statistics for the full sample (N = 1051) and the sample for which we were able to obtain census data (N = 1000). Census characteristics are presented in Table 2. The characteristics of the two samples are not substantively different. Most patients were healthy with few comorbid conditions. Approximately 40% of patients had a mastectomy, and 44% of these had reconstruction. The most common histologic subtype was cribriform (36%) followed by comedo (31%). Multifocal disease was present in 34% of the patients.

Lymph node dissection was performed in 375 patients (35% of the full sample). Of these, 43 (12%) were sentinel node dissections. Patients who had a mastectomy were more likely to have lymph node assessment performed than were those treated with breast conserving surgery (78% versus 8%, $p < 0.001$) regardless of whether or not the patient had breast reconstruction. Women who were younger at diagnosis were also more likely to have a lymph node dissection (54% of women under 40 years compared to 30% of women over age 65 years, $p = 0.002$). Other factors associated with higher likelihood of lymph node assessment were diagnosis earlier in the study period ($p < 0.001$), higher nuclear grade ($p < 0.001$), multifocal disease ($p = 0.03$), and greater mammographic extent of disease ($p < 0.001$). Race, presence of calcifications (not shown), and comorbidity were not associated with lymph node dissection in bivariate analysis. Woman who had lymph node dissection were more likely to have a mastectomy with autogenous reconstruction in the later years and were more likely to have private insurance in the later years (results not shown).

Table 3 presents results from the multivariate model. We present estimates from the census sample (N = 1000) because of the similarity in the results with those from the full sample (N = 1051). We also found our results to be robust to different specifications in calendar time, census-based SES measures, and the inclusion of insurance by time interactions. In addition to the parameter estimates shown in Table 3, our models included controls for comorbid conditions and disease characteristics (including all measures shown in Table 1), findings of which were substantively consistent with the literature. The model fit well. Calibration (HL = 12.250, $p = 0.834$) and discrimination (C = 0.942 for both models) were excellent, and the pseudo R-squared statistic was high at 0.555.

Surgical Treatment—Surgical treatment was strongly predictive of lymph node dissection. Women who had breast conserving surgery were far less likely to undergo lymph node dissection than woman who had mastectomy and no reconstruction (OR = 0.018, $p < 0.001$). Women who had mastectomy with autogenous reconstruction (OR = 3.711, $p = 0.016$) were far more likely to have lymph node dissection than women who had mastectomy without reconstruction. Mastectomy with unknown reconstruction was strongly associated with higher rates of lymph node dissection, but these estimates were imprecise because of the small sample size (N = 8 women who had unknown reconstruction in the estimation sample). The results were unchanged when we re-estimated the model dropping these 8 observations. Specifications in which we interacted surgical treatment and calendar time revealed no evidence that the relative differences by treatment varied over the study period. Very few cases of mastectomy with autogenous reconstruction were done before 1994, however, limiting the interpretability of these results.

Socioeconomic Factors and Demographics—Neither age nor race was found to be a statistically significant or substantively important predictor of lymph node dissection. We estimated many specifications using a variety of census-based SES measures, with and without census-based race measures and with and without individual-level race measures. In none of these specifications was race a statistically significant predictor, and no other

covariate values were substantively affected. Census-based income, however, as measured by Percent Below Poverty, predicted lower rates of lymph node dissection (OR = 0.98, $p = 0.04$). Exclusion of income did not change the estimates of either the individual-level or census level race variables. The addition of other income and wealth related census-tract variables (such as home ownership and educational attainment) generated similar patterns of results to percent below poverty but were imprecisely estimated because of the high correlation among these variables.

Calendar Time—One of our *key hypotheses* was that rates of lymph node dissection would decrease over time as best practice information diffused. Our results strongly confirm this. Our specification of a second order polynomial in calendar year (Year and Year squared) uncovered a strong and statistically significant downward trend in lymph node dissection during the study period (as evident in Figures 1 and 2).

Calendar Time and Site Interactions—A second *key hypothesis* was that lymph node dissection rates at different sites would converge over time as best practice information diffused. Again, our results strongly support this finding. Although we found large and statistically significant differences in the rates of lymph node dissection between sites early in the study period (OR = 3.574, $p = 0.026$), the differences converged to nearly zero in the late 1990's (OR = 1.192, $p = 0.642$). Figure 1 presents the results graphically and shows that the rates of lymph node dissection decrease rapidly with time, reaching a low in the late 1990's and that the differences between sites converged dramatically around 1995. Different specifications of time interactions by site generated substantively similar conclusions, with different rates of decrease by site causing a divergence in lymph node dissection rates that subsequently diminished.

Calendar Time and Insurance Status Interactions—A third key hypothesis was that differences in lymph node dissection rates would be driven by the generosity of insurance but that differences in rates of lymph node dissection among women with different types of insurance would diminish with time. Again, we found evidence that supports the hypothesis. We found no evidence that women with no insurance or Medicaid insurance had rates of lymph node dissection that differed from women with Medicare only (our referent group). More generous insurance, however, was strongly associated with increased rates of lymph node dissection early in the study period. In particular, women who had private insurance (OR = 3.363, $p = 0.003$) or Medicare with supplemental private insurance (OR = 4.246, $p = 0.011$) had higher rates of lymph node dissection than women with Medicare only early in the study period. We found that these very large differences began to diminish in the mid 1990's, and that there were only small and not statistically significant differences across these insurance types by the late 1990's (OR = 1.272, $p = 0.552$ and OR = 0.625, $p = 0.449$ for private and Medicare with supplemental insurance, respectively). Estimates of our overlapping polynomial parameters suggest that the convergence between Medicare (the referent group) and Medicare plus private insurance was centered in 1997 and occurred largely over a 4-year period from 1995 through 1999 ($\sigma = 1.2$). Our estimates indicate that the convergence between Medicare and private insurance was centered in 1995 and occurred over a slightly longer period from roughly 1992 through 1998 ($\sigma = 2$). The substantive consequence of the overlapping polynomial estimates is illustrated in Figure 2, which presents our estimates graphically. The Figure shows the strong downward trend in lymph node dissection for each of the insurance types during the study period. It also shows the large differences in rates of lymph node dissection by insurance in the late 1980's and early 1990's and the paths of convergence of these rates in the late 1990's. By 1998, the remaining differences in lymph node dissection rates by insurance status were neither large nor statistically significant.

Discussion

In this large sample of women diagnosed with DCIS between 1985 and 2000, lymph node dissection rates varied significantly according to site of treatment and type of insurance after controlling for clinical, pathological, treatment, and demographic characteristics. The decrease in lymph node dissection in women with DCIS in our sample is consistent with that seen in larger samples of patients in Surveillance, Epidemiology, and End Results registries.^{28, 29} As we had hypothesized, the passage of time was associated with a decrease in the site and insurance variation in lymph node dissection rates, variation that was no longer significant by the end of the study period. This finding may reflect the diffusion of information about best practices facilitated by published practice guidelines¹² and expert consensus.³⁰ In addition, provider experience or patient knowledge may drive the divergent treatment rates together over time although it is unlikely that patients have much to do with the changing rates unless their information is coming from an available source that aggregates the information (such as guidelines).

The higher likelihood of lymph node dissection among women with Medicare plus private insurance compared to those with Medicare alone supports our hypothesis that generosity of insurance plays a role in surgeons' treatment decisions. Although we do not actually know the size of the financial incentives because we do not know the differences in reimbursements and copayments, those with Medicare and supplemental private insurance have more comprehensive coverage than those with Medicare only. Private insurance (under age 65 with no Medicare) is heterogenous, being very generous for some and leaving others with very poor protection. This results in different expected reimbursement for physicians and different out-of-pocket costs for patients. It is also possible that patient demand for lymph node dissection is greater when insurance blunts the cost constraints to patients. While surgeons may be more inclined to perform procedures with higher fees (all things being equal), the choice of procedure is likely to be influenced by how much a patient is willing to pay.

Our findings are consistent with previous research. For example, in a study of 1,787 women with breast cancer, decreasing Medicare mastectomy fees were associated with increasing use of breast conserving surgery.⁸ Further support for the role of payment incentives in surgical treatment of breast cancer comes from a study demonstrating that area Medicare fees were the strongest predictors of "propensity" (surgical treatment behavior) in scenarios presented to breast surgeons.³¹

In general, greater generosity of insurance, such as might be the case in patients with Medicare plus supplemental insurance compared with Medicare alone, is considered a favorable condition, increasing access to health care and, in one recent study, access to clinical trials among Medicare patients.³² Little attention has been given to the role of type of insurance and overtreatment. Over 20 years ago, Ware and colleagues advised against assuming that the relationship between generosity of health insurance and health status is always positive: "More generous health insurance may be a two-edged sword."³³ Our findings support this concern. While one would not argue that receipt of lymph node dissection negatively affects survival, unnecessary lymph node surgery adversely affects quality of life^{34, 35} and contributes to the cost of medical care.^{36–38}

In conclusion, we have demonstrated that rates of a discretionary surgical procedure differ according to non-clinical factors, such as treatment site and type of insurance, and that such unwarranted differences decrease over time with diminishing uncertainty and, perhaps, through diffusion of clinical guidelines. If similar analyses were to be repeated in contemporary patient samples for receipt of sentinel node dissection, a procedure considered less invasive because of removal of fewer lymph nodes, our findings of site and insurance

variations may very well be seen. The lack of consensus regarding the need for sentinel node dissection in DCIS19, 39–43 and non-clinical variation in receipt of sentinel node dissection in women with invasive breast cancer^{44, 45} would be expected to lead to a recapitulation of similar variation that may in time diminish or resolve.

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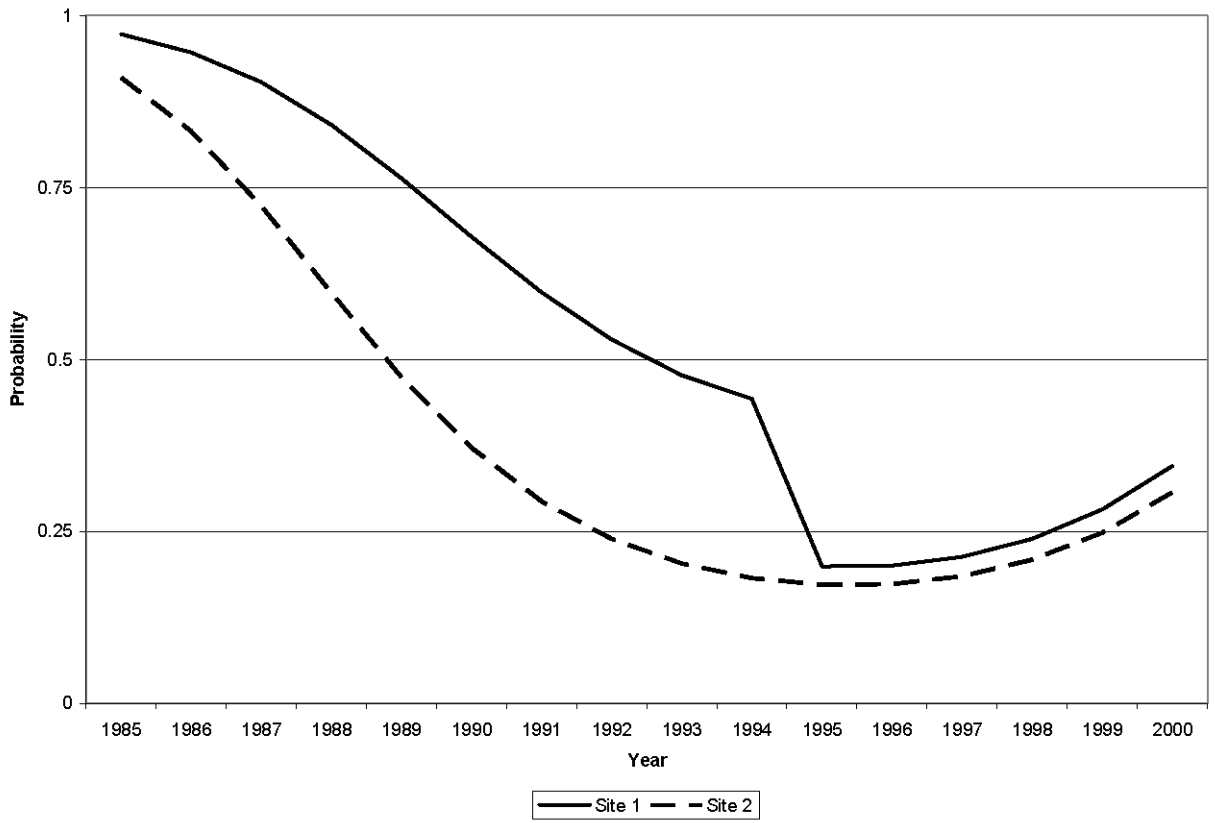


Figure 1.
Estimated lymph node dissection rates by site over time, 1985 – 2000.

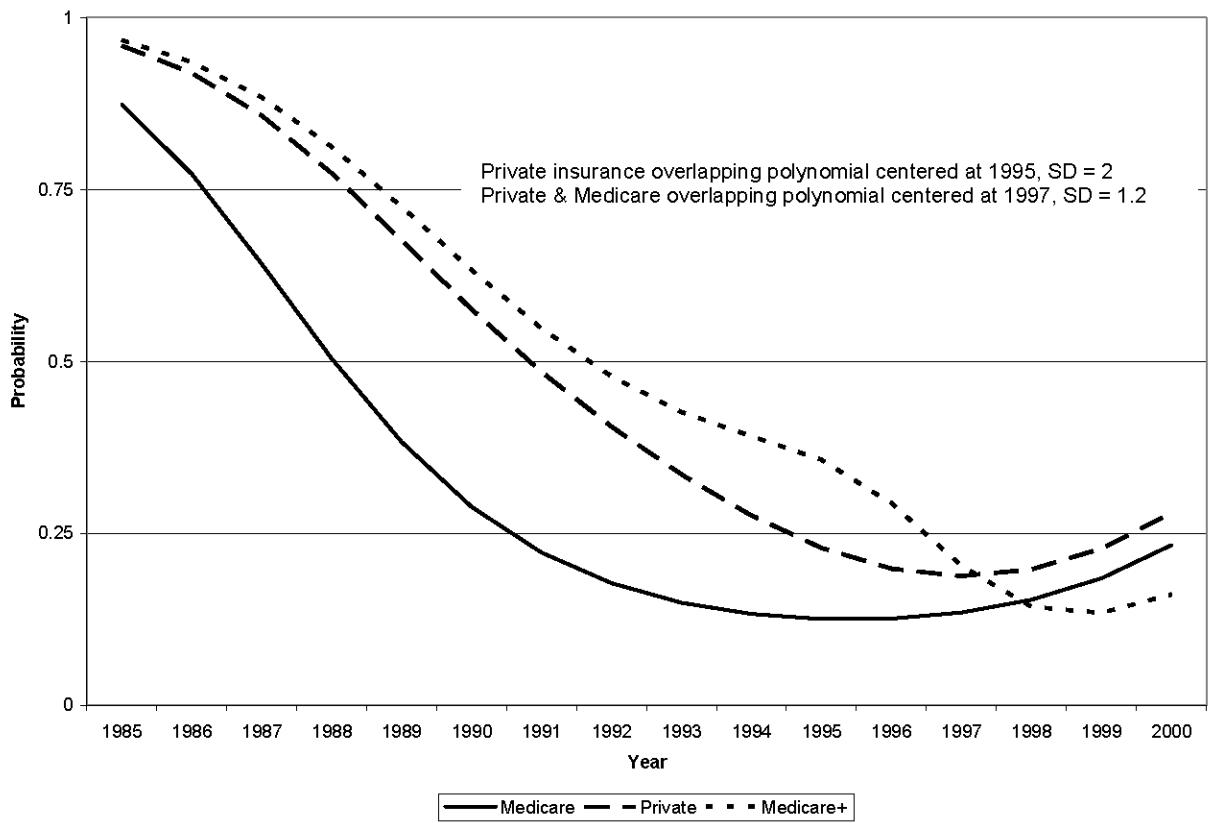
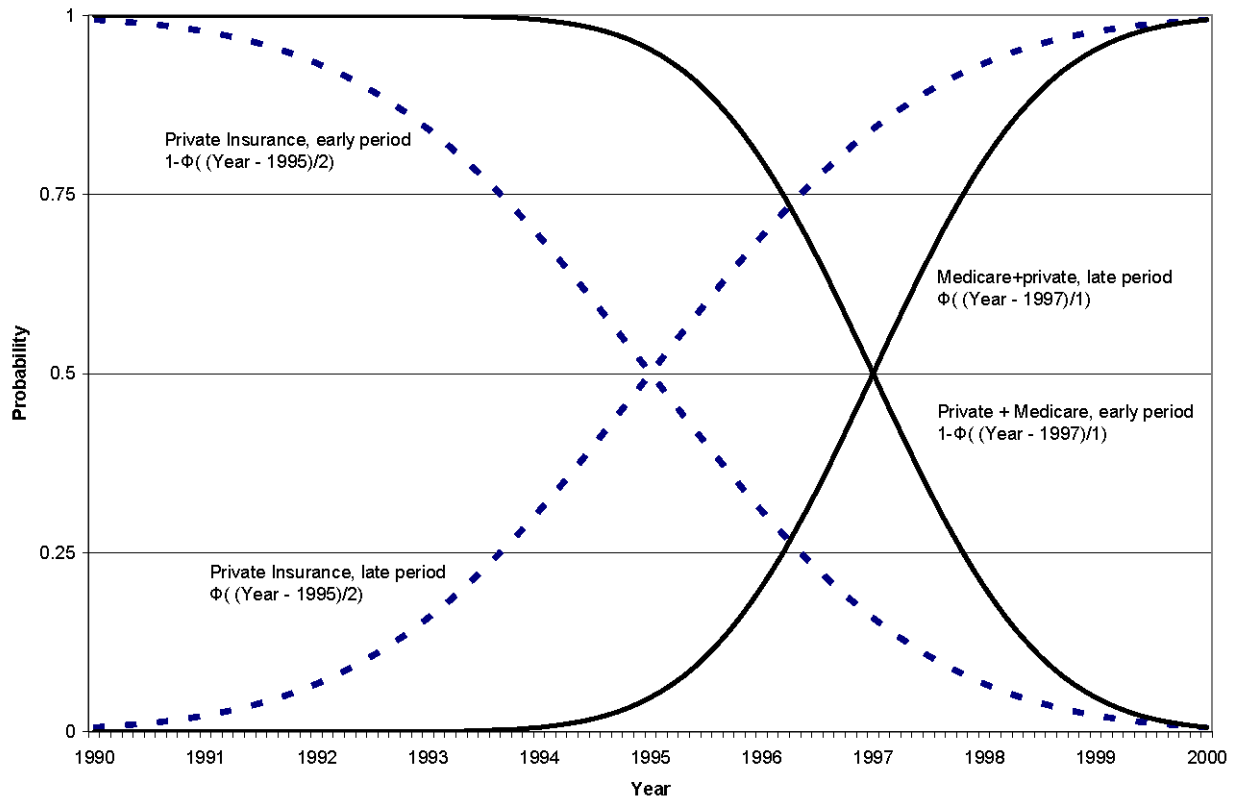


Figure 2.
Estimated lymph node dissection rates by insurance type over time, 1985 – 2000.



Appendix.
Examples of overlapping polynomials

Table 1

Sample Characteristics, Individual-Level Measures

	Percent with Characteristic												P	
	Full Sample				Census Sample				Without LND					
	N	%	N	%	N	%	N	%	N	%	N	%		
Total	1051		1000		349	34.9	651	65.1						
Age (years)														
Less than 40	54	5.1	49	4.9	27	7.7	22	3.4						<0.001
40 to 54	241	22.9	230	23.0	92	26.4	138	21.2						
55 to 64	376	35.8	357	35.7	121	34.7	236	36.3						
65+	380	36.2	364	36.4	109	31.2	255	39.2						
Calendar Year of Diagnosis														
Before 1987	35	3.3	34	3.4	29	8.3	5	0.8						<0.001
1988–1991	107	10.2	104	10.4	63	18.1	41	6.3						
1992–1993	124	11.8	117	11.7	51	14.6	66	10.1						
1994–1995	162	15.4	152	15.2	43	12.3	109	16.7						
1996–1998	451	42.9	426	42.6	121	34.7	305	46.9						
1999–2000	172	16.4	167	16.7	42	12.0	125	19.2						
Insurance														
Medicare	324	30.8	307	30.7	90	25.8	217	33.3						<0.001
Private	537	51.1	508	50.8	194	55.6	314	48.2						
Medicare + Private	106	10.1	102	10.2	37	10.6	65	10.0						
Medicaid/uninsured/other	18	1.7	18	1.8	5	1.4	13	2.0						
Unknown	66	6.3	65	6.5	23	6.6	42	6.5						
Comorbidity[†]														
Osteoarthritis	191	18.2	178	17.8	59	16.9	119	18.3						0.59
Asthma	67	6.4	65	6.5	22	6.3	43	6.6						0.85
Depression	76	7.2	71	7.1	20	5.7	51	7.8						0.22
Diabetes with insulin	49	4.7	49	4.9	13	3.7	36	5.5						0.21

	Percent with Characteristic						Census Sample						p
	Full Sample			Census Sample			With LND			Without LND			
	N	%		N	%		N	%		N	%		
Diabetes no insulin	24	2.3	23	2.3	6	1.7	17	2.6	0.37				
Congestive heart failure	29	2.8	29	2.9	6	1.7	23	3.5	0.10				
Hypertension	380	36.2	362	36.2	116	33.2	246	37.8	0.15				
Smoking	171	16.3	162	16.2	49	14.0	113	17.4	0.17				
Thyroid	121	11.5	119	11.9	39	11.2	80	12.3	0.6				
Race									<0.001				
White	835	79.4	794	79.4	283	81.1	511	78.5					
Black	158	15.0	151	15.1	47	13.5	104	16.0					
Asian	10	1.0	9	0.9	5	1.4	4	0.6					
Unknown	47	4.5	46	4.6	14	4.0	32	4.9					
Histologic Subtype													
Comedo	333	31.7	314	31.4	138	39.5	176	27.0	<0.001				
Cribiform	382	36.3	364	36.4	105	30.1	259	39.8	<0.001				
Apocrine	10	1.0	9	0.9	4	1.1	5	0.8	0.55				
Other	23	2.2	22	2.2	12	3.4	10	1.5	0.05				
Unknown	50	4.8	50	5.0	10	2.9	40	6.1	0.02				
Papillary	44	4.2	43	4.3	9	2.6	34	5.2	0.05				
Micropapillary	74	7.0	70	7.0	17	4.9	53	8.1	0.05				
Nuclear Grade									<0.001				
Low	139	13.2	131	13.1	19	5.4	112	17.2					
Intermediate	199	18.9	196	19.6	47	13.5	149	22.9					
High	436	41.5	409	40.9	162	46.4	247	37.9					
Unknown	277	26.4	264	26.4	121	34.7	143	22.0					
Multifocal	358	34.1	342	34.2	135	38.7	207	31.8	0.03				
Surgical Treatment									<0.001				
Mastectomy	414	39.4	395	39.5	306	87.7	89	13.7					

	Percent with Characteristic						Census Sample						p
	Full Sample			Census Sample			With LND			Without LND			
	N	%		N	%		N	%		N	%		
Reconstruction													< 0.001
Implant	65	6.2		62	6.2		42	13.7		20	22.5		< 0.001
Autogenous	97	9.2		95	9.5		77	22.1		18	2.8		< 0.001
None	233	22.2		220	22.0		182	59.5		38	42.7		< 0.001
Unknown	7	0.7		6	0.6		5	1.6		1	1.1		0.01
Breast Conserving Surgery	637	60.6		605	60.5		43	12.3		562	86.3		
Site Dummy	638	60.7		604	60.4		246	70.5		358	55.0		< 0.001

[†] Some patients have more than one comorbid condition

Table 2

Summary Statistics, Census-Level Measures

	Total Sample		No LND		LND		p value
	Mean	SD	Mean	SD	Mean	SD	
Percent black	17.1%	30.7%	18.7%	32.5%	14.0%	27.0%	0.02
% Not speaking English	1.7%	3.1%	1.7%	2.9%	1.9%	3.5%	0.38
Education							
Percent with less than a HS education	15.1%	11.8%	15.1%	11.5%	15.1%	12.3%	0.96
Percent with at least a HS diploma	28.0%	10.6%	27.9%	11.1%	28.3%	9.8%	0.60
Median household income	\$53,035	\$24,570	\$52,794	\$25,017	\$53,485	\$23,745	0.67
Percent below poverty	9.6%	11.2%	10.0%	11.3%	8.9%	11.0%	0.14
Percent home owners	75.4%	24.1%	74.5%	24.6%	76.9%	23.1%	0.13

HS = high school

Table 3

Multivariate analyses

Census Sample (N = 1000)			
Dependent Variable = lymph node dissection	OR	SE	p
Surgical Treatment			
Breast Conserving Surgery	0.02	0.01	0.000
Mastectomy/Reconstruction			
Referent	1.00		
Implant	0.58	0.21	0.13
Autogenous	3.71	2.02	0.02
Unknown	3.67	3.85	0.22
Census Tract			
Percent below poverty	0.98	0.009	0.04
Site * Treatment Interactions			
BCS	0.20	0.12	0.007
Mastectomy with autogenous reconstruction	0.12	0.08	0.003
Calendar Time			
Year	0.44	0.08	0.000
Year ²	1.04	0.01	0.000
Site			
Site * Year < 1995 (early period)	3.57	2.05	0.026
Site * Year > 1994 (late period)	1.19	0.45	0.64
Insurance Coverage (Referent = Medicare)			
Private* $\Phi(1 - ((t-1995)/2))$ (early period)	3.36	1.36	0.003
Private* $\Phi((t-1995)/2)$ (late period)	1.27	0.51	0.55
(Private+Medicare)* $(1 - \Phi((t-1997)/1.2))$ (early period)	4.25	2.43	0.01
(Private+Medicare)* $\Phi((t-1997)/1.2)$ (late period)	0.63	0.39	0.45
Medicaid/uninsured/other	0.90	0.76	0.90
Unknown	1.18	0.59	0.75
Goodness of Fit Statistics			
Wald test of Significance of Model	311.480		0.000
Pseudo R²	0.555		
C-Stat	0.942		
Hosmer-Lemeshow, 20 groups, chi²(18)	12.250		0.834