Physician Agency and Competition: Evidence from a Major Change to Medicare Chemotherapy Reimbursement Policy*

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Abstract

We investigate the role of physician agency and competition in determining health care supply and patient outcomes. A 2005 change to Medicare fees had a large, negative impact on physician profit margins for providing chemotherapy treatment. In response to these cuts, physicians increased their provision of chemotherapy and changed the mix of chemotherapy drugs they administered. The increase in treatment improved patient survival. These changes were larger in states that experienced larger decreases in physician profit margins. Finally while physician response was larger in more competitive markets, survival improvements were larger in less competitive markets.

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Understanding how physicians respond to financial incentives is a key health care policy issue. In the presence of asymmetric information, physicians may distort demand in socially sub-optimal but personally beneficial ways. In the U.S. context, debate over this issue centers on physician-induced demand (PID) – the provision of excessive care in response to financial incentives (McGuire (2000); Chandra et al. (2012)).

As the theoretical literature on PID makes clear, however, the relationship between administered price changes and the volume of physician services is a priori ambiguous and depends on the relative strength of the income and substitution effects (McGuire and Pauly 1991).[1] When the substitution effect dominates, physicians reduce the volume of services in response to a price reduction. When the income effect dominates, the physician's supply curve will have a negative slope such that volume increases as fees fall. The case of extremely large income effects corresponds to a model of "literal target income," where changes in administered prices are fully offset by changes in volume (assuming full offset is feasible).

The empirical literature supports the importance of physician agency, with the majority of papers finding a negative relationship between changes in fees and the supply of medical care, i.e. volume offsets (Rice and Labelle (1989); Rice (1983); Rice (1984); Christensen (1992); Yip (1998); Nguyen and Derrick (1997); Grant (2009); Gruber et al. (1999)).[2] Importantly the existing literature has found no direct evidence that such changes affect patient health, although few studied measure outcomes.[3]

We contribute to the literature on volume offsets in two ways. First, using a very general theoretical model of physician utility, we show that physician agency can distort the provision of services away from the fully informed consumer's optimal level of care in either direction. That is, physicians may provide excessive care or may ration care below the optimal amount. The implication of this theoretical point

[1]This conclusion from McGuire and Pauly (1991) comes out of a model where physicians maximize utility that depends positively on net revenue and leisure and negatively on the amount of demand inducement.

[2]Clemens and Gottlieb (2012); Hurley et al. (1990); Hurley and Labelle (1995); Escarce (1993) are notable exceptions that find a positive relationship between fees and the quantity of health care.

[3]Clemens and Gottlieb (2012) is again an exception. Most evidence is indirect, drawn from the observation, for example, that elective C-sections have more complications than vaginal delivery.

is that a fee cut that generates an income effect that dominates the substitution effect can remedy an underprovision of care. Furthermore, we show that although the response to a payment change should be smaller in less competitive markets, the welfare effects will be larger. The intuition is that in more competitive markets, those patients who need care are more likely to get it and care is at the 'flat of the curve.'

Second, we analyze an important empirical case, a 2005 change to Medicare fees that likely had a large impact on the income of a particular specialist group, oncologists. Specifically, we study a major reform to Medicare reimbursement policy for physician-administered (Part B) drugs, primarily injectable and intravenous oncology, rheumatology, urology and infectious disease agents. Unlike orally administered drugs, i.e. pills, physicians purchase Part B drugs from wholesalers, administer them in their clinics, and are reimbursed for them directly by payers. The 2005 reform, which replaced a reimbursement system based on list prices with one based on national average transaction prices, sought to remedy well known Medicare overpayments for Part B drugs. For chemotherapy, reimbursements averaged 29% above transaction prices in 1997 (Office of Inspector General (1997a)) and was 22% above by 2004, even after an initial fee reduction took effect (Government Accountability Office (2004)). Beginning in 2005, margins were administratively set at 6% above the national average transaction price for each drug. Thus, margins fell by nearly a factor of five, although this reduction was not uniform across drugs. Since chemotherapy accounted for over three-quarters of oncologist practice revenues in 2005 (Akscin and Barr (2007)), the drug mark-up accounted for more than half of their reported net income (Butcher (2008)), and because Medicare is the primary payer for many cancers, the reform almost assuredly represented a large negative shock to oncologist income.[4]

Our empirical analysis centers on comparing Medicare beneficiaries diagnosed with lung cancer in the few months before to the few months after the reform took effect. After demonstrating that the observable characteristics of lung cancer patients were smooth across the payment change, we show that those diagnosed just after relative

[4]Oncologists maintained that Medicare fees for other services, e.g., evaluation and management and drug administration, did not cover their expenses. However, GAO has found that these payments exceed expenses by 4% on average, which is similar to the mark-up received by other specialists (Government Accountability Office (2001)).

to just before the change were about 10% more likely to be treated with chemotherapy within 30 and 90 days of diagnosis. The change lines up with the timing of the 2005 reform and is specific to the physician's office setting in contrast to the hospital outpatient setting, where the payment change did not take effect until 2006 Barlas (2011). Furthermore, we show the mix of drugs used to treat patients changed in ways predicted by economic theory: drugs that lost the most margin were used less frequently among the chemotherapy-treated population and conversely, expensive drugs favored by the reform's 6% margin on all drugs, were used more often than previously.

Further, we find that the treatment changes improved survival: the likelihood of death at 3, 6, and 9 months from diagnosis declined by 2–3%, suggesting that, consistent with the clinical literature, lung cancer may be undertreated in the elderly (Davidoff et al. (2010); Booth C. M. et al. (2010)). Indeed, despite similar relative treatment increases, the improvement in survival was 1.5 to 3 times larger in above vs. below median Medicare age patients, those most likely to be undertreated. We exploit geographic heterogeneity in treatment responses (Jacobson et al. (2011)) to show that survival improved more in states with larger treatment increases. We also exploit pre-reform prescribing patterns to show that both the treatment increase and survival improvements were larger in areas that experienced a larger decrease in profit margins. As best as we know, this is among the first evidence that physician agency can affect health outcomes, not just treatment. Finally, consistent with our model, we find that, although the treatment change was more muted in less competitive counties (measured by the HHI or the number of patients per provider), survival improvements were larger. These results suggest that the competitive environment affects both the size of physician response and patient welfare. To our knowledge, physician market power has not been studied in the context of changes in administered prices, perhaps because it was assumed to not be relevant to administered price regimes.

1 Institutional Background

Although Medicare did not offer a drug benefit for oral drugs (pills) until 2006, Medicare Part B, which covers physician services, has from inception covered physician-administered drugs such as IV chemotherapy, anti-nausea, and pain medicine used in

cancer treatment. Rather than writing a prescription for these drugs that the patient fills at a pharmacy, the physician injects or infuses these drugs in an office or clinic and then bills Medicare for them.[5] Since at least the mid-1990s, it was well understood that Medicare overpaid for many chemotherapy and other Part B drugs (Office of Inspector General (1997b); Office of Inspector General (1997a)). The Medicare Prescription Drug, Improvement and Modernization Act (MMA) of 2003 aimed to resolve the overpayment issue and reduce Medicare spending by redesigning the way Medicare reimburses physician-administered (Part B) drugs.

Prior to the reform, Medicare reimbursed physicians and outpatient hospital clinics for these drugs as a percentage of the average wholesale price (AWP), a list price, published in several catalogues, that reflects neither wholesale nor average prices (Berndt and Newhouse (2012)).[6] From 1998 to 2003, reimbursements were 95% of AWP. To smooth the transition to a new payment system, the MMA reduced reimbursements to 85% of AWP in 2004. In principle manufacturers could have changed list prices between 2003 and 2004 (or in other years); in practice, such changes were rare. Most importantly for our purposes, the AWP exceeded physician acquisition costs for many drugs from 13% to 34% (Government Accountability Office (2004)). Some agents, such as paclitaxel, a chemotherapy drug commonly used to treat lung, breast, and ovarian cancer, were reimbursed at prices that vastly exceeded acquisition costs (Government Accountability Office (2004)). The reform replaced the AWP-based system with a new average sales price (ASP) payment system, whereby Part B drugs are reimbursed based on the national average of manufacturers' sales prices, including rebates, from two quarters prior plus a 6% mark-up.[7] In 2005 the reform affected physician's offices, the primary setting for chemotherapy treatment, and in 2006 the outpatient hospital setting, the other main setting for chemotherapy.

[5]In contrast, some private insurers deliver the drug to the physician's office or clinic and just pay the physician a fee for administration.

[6]In the trade, AWP is sometimes referred to as "Ain't What's Paid."

[7]The ASP calculation does not take into account 340(B) discounts for Medicaid patients. Since about 2009 it is believed that hospitals began leveraging these discounts to purchase oncology drugs for all their patients and thereby increase the mark-ups they receive on oncology drugs administered to Medicare and commercially insured patients. This is a relatively recent practice that is unlikely to have occurred during our study period.

When this policy took effect on January 1, 2005, the profit margins for many chemotherapy drugs were reduced substantially. Fees for chemotherapy administration were increased in 2004 but reduced again in 2005, though the reduction did not fully offset the 2004 increase (Government Accountability Office (2004)). Importantly, the administration fee changes were small in dollar terms (on the order of $50 to $100 per session) compared with the 2004 to 2005 drug reimbursement change, which could total hundreds and even thousands of dollars for a monthly dose of some commonly infused drugs.[8]

Two health policy papers have evaluated the claim by community oncologists that the policy change would make it too expensive to treat Medicare beneficiaries, forcing them to shift some patients to hospital settings where they would face unnecessary treatment delays (Community Oncology Alliance (2006)). The papers find to the contrary; wait times, travel distance and treatment setting (outpatient hospital or physician office) did not change in response to the payment change (Friedman J. Y. et al. (2007); Shea et al. (2008)). Neither paper evaluates treatment outcomes. Moreover, both studies – one, a web-survey of a convenience sample of patients undergoing chemotherapy (Friedman J. Y. et al. (2007)) and the other a 5% sample of claims from beneficiaries with leukemia, lymphoma or breast, colorectal, lung cancer receiving chemotherapy – condition on treatment, a potentially important margin of adjustment.

Indeed, our prior work shows that the likelihood of chemotherapy treatment for lung cancer patients changed in response to the reform (Jacobson et al. (2010)). Our findings are consistent with analyses of aggregate chemotherapy billing data by the Medicare Payment Advisory Committee (MedPAC (2006)). Other work finds important extensive margin responses. For example, in prostate cancer treatment, androgen-deprivation therapy (ADT) or medical castration, which was subject to AWP-reimbursement, declined while surgical castration, which did not face a payment change, increased (Weight et al. (2007)). Although analysis of ADT suggests that the decline was largest among those for whom the benefits were unclear (Shainian et al. (2010)), these welfare implications may not generalize to other cancers.[9]

[8]Many regimens are given in 3-week cycles, meaning patients receive 1.33 treatments per month.

[9]Colla et al. (2012) finds that in response to the reform chemotherapy treatment declined in the

An obvious question is whether these treatment changes were driven by physicians or patients; since Medicare beneficiaries face 20% coinsurance for Part B services, a decline in drug prices could have increased patient demand. In practice, however, about 90% of beneficiaries have supplemental insurance that pays this coinsurance (Kaiser Family Foundation (2010)). And anecdotal evidence suggests that oncologists were less likely to collect the coinsurance from the remaining patients when reimbursements were based on AWP, implying that out-of-pocket spending increased for some beneficiaries even as average drug costs declined for Medicare (Mullen (2007)). On balance, any effects from patient demand should be negligible.

2 Health Care Supply

In this section we present a model of physician behavior in markets where physicians care about both their private utility as well as the welfare of their patients.

Let there be J physicians operating in a market $k \in \{1, K\}$ that serves a continuum of patients of measure one. A physician $j \in \{1, J\}$ has local monopoly power over a fraction η_j of patients in market k and can provide each patient a single unit of service q, which yields patient i a benefit b distributed over $(-\infty, \infty)$ such that b(q) is a continuous, decreasing convex function. In addition, the distribution of patients each physician sees is identical to each other and to the overall distribution of types in the market.

The physician earns a fixed price p and pays a fixed cost c for each unit of service provided. Each unit of care requires a unit of physician time such that e(q), the disutility of physician effort, is an increasing and convex function with $e(0) = 0$. Physician utility is assumed to be positive in net income with diminishing marginal returns. In addition, physicians are assumed to care directly about patient welfare. Assuming additive separability, the physician utility function can then be written as

$$U(q) = V(\pi) - e(q) + \alpha\eta \int_0^{q/\eta} b(x)dx, \tag{1}$$

last 14 days of life.

where $\pi = (p - c)q$ and α represents the weight physicians place on patient benefit. Thus, a physician trades off her private utility from income with the disutility from of effort as well as her concern for patient welfare.

For notational simplicity, we define the profit margin $m = (p - c)$. Then the first order condition for the physician is given by

$$mV_\pi - e_q + \alpha b(q/\eta) = 0. \tag{2}$$

Note that while concern for patient welfare ($\alpha b(q/\eta)$) pushes physicians towards providing what, from the patient's perspective, is the optimal level of care q^P,[10] in general physicians may either under or over provide care relative to this optimum.

To determine physician response to profit margin (m) changes, we take the derivative of equation 2 with respect to m. Rearranging, we get the following relationship:

$$q_m = \frac{V_\pi + V_{\pi\pi}\pi}{e_{qq} + \frac{\alpha}{\eta}b'}. \tag{3}$$

Note that in general q_m can be positive or negative. That is, it cannot be determined ex-ante whether an increase in profitability leads to an increase or decrease in service provision. This ambiguity of changes in m on the quantity of service provided is a standard characteristic of PID (see for instance McGuire and Pauly (1991) and McGuire (2000)), and in the case of a fee cut is driven by the tradeoff between decreasing marginal returns (the substitution effect) and the increasing marginal returns to income (the income effect). So if decreased returns to effort dominates, a cut in margins leads to a reduction in effort (i.e. services provided). But when the income effect dominates, a decrease in margins can generate a negatively sloped supply curve. In this case, a fee cut can lead to an increase in physician effort. Thus the change in the marginal utility of income (i.e. the relative magnitudes of V_π and $V_{\pi\pi}$) provides some guidance as to the sign of the response. When the magnitude of the income effect $V_{\pi\pi}$ is small, then the substitution effect dominates and physicians display a positive elasticity. But when $V_{\pi\pi}$ is large, it reduces the impact of the substitution effect, and if large enough, can lead to a negative elasticity of health care. In addition, regardless

[10] In the limit $\alpha \to \infty$, $q^* \to q^P$.

of the sign of q_m, physician response is attenuated by the weight physician's place on patient welfare α, and in the limit $\alpha \to \infty$, $q_m \to 0$.

As implied by the name, the literature frames physician-induced demand (PID) as the inducement of excess demand, and not the idea that physicians would "hold down" demand. But as the FOC in equation 2 makes clear, whether the equilibrium quantity q^* is above or below the patient bliss point q^P is ambiguous: physician self-interest can distort the provision of services away from q^P in either direction. In principle, though not in practice, PID could be defined as influencing patient demand away from q^P (either through under or over provision). Influencing patient demand towards q^P could be defined as "useful agency." Importantly, whether changes in care are helpful or harmful to patient welfare depends not only on the direction of the change but also the original level of service relative to q^P.

While the idea of PID is complex,[11] for the purposes of this paper we define it based on the concept of distortions away from the care that would be demanded by the mythical, fully informed patient. Specifically we refer to the provision of services to a patient who receives a non-positive benefit (i.e. $b \leq 0$) as physician induced demand (PID). If, on the other hand, the patient would have received a net positive benefit from a service, we call such a distortion physician rationing (PR).

2.1 Market concentration and physician agency

Proposition 1 *For small $\alpha \approx 0$, if there is an internal solution to the FOC in equation 2, the magnitude of the elasticity of market response to changes in profit margin will be negatively correlated with market concentration.*

Proof. See Appendix. ■

Proposition 2 *For small $\alpha \approx 0$, the magnitude of the change in patient benefit in a market will be positively correlated with market concentration.*

Proof. See Appendix. ■

[11] See McGuire (2000) for a detailed discussion of this issue.

2.2 Summary of predictions

In equilibrium, physicians may under or over provide care relative to what would be demanded if all consumers were perfectly informed (q^P). In response to changes in profit margins m, the direction of physician response is ambiguous and determined by the tradeoff between increasing marginal returns and a wealth effect. Additionally, the sign of any change in patient welfare cannot be determined by the sign of the change in the amount of care provided but depends also on the initial level of care.

In terms of market response to changes in m, the magnitude of the treatment change (e.g., the share of patients receiving care) is negatively correlated with market concentration. In contrast, changes in average welfare are positively correlated with market concentration. While the relationship between market concentration and physician response is driven mostly by the mean physician market share $\bar{\eta}_k = \frac{1}{J}\Sigma_{j=1}^{J}\eta_j$ (e.g. the physician to population ratio), the welfare response is driven by both the mean and *distribution* of physician market shares η_j. The mean physician market share determines how many additional patients get treated, the distribution of physician market shares determine *who* gets treated. In short, although the treatment response to a payment change will be larger in more competitive markets, the welfare effects will be smaller.

3 Data and Methods

To study the impact of physician profit margin changes on health care supply and patient well being, we analyze Medicare claims for beneficiaries who had at least one claim in a physician's office or an outpatient-hospital setting with a lung cancer diagnosis (ICD-9 162.0-162.9) between 2003 and 2005 (N=878,923). For these beneficiaries, we have all claims for physicians' services, hospital outpatient care, durable medical equipment, short-stay, long-stay and skilled nursing facility services, and hospice care for 2002 to 2006.[12] These claims capture only Medicare fee-for-service, the care affected by the reform. Although in principle care in Medicare Advantage (MA)

[12]These data come from the Medicare Carrier, Outpatient, Durable Medical Equipment (DME), Medicare Provider Analysis and Review (MedPAR) and Hospice files.

plans, Medicare-approved private health plans that are paid a lump-sub on a risk-adjusted per-enrollee per month basis, could serve as a control, there is no comparable source of treatment data for this population.[13] Dates of death for our analytic cohort (defined below) are from the Medicare Vital Status File through 2012.

A claim with a lung cancer diagnosis may not indicate cancer but could capture a miscoded diagnosis code or evaluation for a suspected cancer that is subsequently deemed benign. To isolate lung cancer cases, we restrict to Medicare beneficiaries with two or more non-institutional (i.e., physician, durable medical equipment or outpatient hospital) claims separated by at least 28 days but no more than 365 days or one institutional (short-stay, long-stay, skilled nursing facility or hospice) claim with a lung cancer diagnosis. This is a common approach to defining cancer in claims data that has been shown to have high sensitivity (i.e., identify a high proportion of actual cancer cases) and very high specificity (i.e., exclude beneficiaries without cancer) in Medicare claims data (e.g., see Warren et al. (1999); Ramsey et al. (2009); Vera-Llonch et al. (2011)). To isolate lung cancer treatments, we exclude beneficiaries with more than one confirmed primary cancer. We use the first claim with a lung cancer diagnosis to date the onset of disease. To ensure new diagnoses and thus treatment trajectories that were not influenced by prior clinical decisions, we restrict to beneficiaries with no cancer-related claims history in the 12 prior months.[14]

We eliminate the roughly 5% of beneficiaries in the fee-for-service claims who were enrolled in a Medicare Advantage plan at some point during the study period and thus lacked complete treatment and billing information. The lung cancer cohort included 216,119 beneficiaries enrolled in Medicare Parts A and B who were diagnosed between January 2003 and November 2005. We further restrict the primary analytic cohort to the 132,768 beneficiaries diagnosed between February 2004 and November 2005 to capture 11 months of data on either side of the payment change.

Chemotherapy treatment is identified using relevant diagnosis and billing codes

[13] Although national MA data are unavailable, we had access to Kaiser Northern California's MA data to use as a control. However, the FFS treatment response to the MMA in Northern California was small, making this comparison less relevant. We do not have a good explanation for geographic variation in the response, although such variation is widespread in Medicare (Skinner (2012)).

[14] Patients for whom we could not verify a 12-month cancer free period are eliminated.

(Warren et al. (2002)).[15] Beneficiary date of death is drawn from the Medicare Vital Status File, which we had updated through June 30, 2012.

3.1 Analytic Approach

To assess whether the January 2005 payment change plausibly affected treatment and health outcomes, we take both a graphical and regression-based approach that initially analyzes patients by month of diagnosis relative to the reform. To begin, we plot estimated month-relative-to-reform (January 2005) fixed effects along with 95 percent confidence intervals from a regression of the likelihood a newly diagnosed lung cancer patient received any chemotherapy treatment, chemotherapy treatment by setting (e.g., physician's office), treatment with specific agents conditional on receiving any chemotherapy, all within 30 days of diagnosis. We choose a 30-day treatment period because it enables us to relatively cleanly define treatment (diagnosed post-reform) and control (diagnosed pre-reform) cohorts.[16] The month-relative-to-reform fixed effects are simply the set of coefficients from the following regression:

$$Y_{ism} = \beta_0 + X_{ism}\gamma + \mu_s + \delta_{m-m^*} + \epsilon_{ism}, \tag{4}$$

where Y_{ism} is the treatment or survival outcome of individual i residing in state s and diagnosed in month m, X_{ism} are a set of patient characteristics including gender, race/ethnicity (7 categories), patient age and its square, pre-cancer comorbidities, as measured by the Deyo-Charlson score (Deyo et al. (1992); Charlson et al. (1987)), an indicator for metastatic disease at diagnosis, defined as 30 days from the date of diagnosis[17]; μ_s are state fixed effects, δ_{m-m^*} are month-relative-to-reform fixed effects

[15] We use ICD-9-CM procedure code 99.25 or diagnosis codes V58.1, V66.2 or V67.2, diagnosis-related group code 410, or HCPC codes 96400-96549, J9000-J9999 or Q0083 through Q0085, C8953-C8955 (2006 only), G0345-G0362 and revenue center codes 0331, 0332 and 0335.

[16] Longer treatment windows capture cohorts diagnosed in the pre-reform period but receiving care in the post-reform period. We show results for these groups in sensitivity checks.

[17] To confirm metastasis we require one institutional claim or two or more non-institutional claims separated by at least 28 and no more than 365 days with a secondary cancer diagnosis [ICD-9 codes: 197-199]. Although this algorithm is likely to miss many beneficiaries with secondary cancer, it does have reasonable specificity (Earle et al. (2002)). Results are not sensitive to excluding this measure.

(e.g. $mx - m^* = -1$ in December 2004 and $+1$ in February 2005) with January 2005 omitted, and ϵ_{ism} is an error term. Plotting these month fixed effects helps verify that the time-series patterns are consistent with a causal impact of the reform. In addition, we use these plots to guide the specification of models estimating the magnitude of the reform's impact. We focus on a window of 11 months prior to and 11 months post reform, which was chosen to balance the number of pre-reform months with the limited post-reform data available to us, although using all 24 months of pre-reform data yields similar results.[18] We show results for shorter windows - 9 months on either side of the reform - to hone in on the policy change.

We use the same basic regression framework and visual approach to analyze both the characteristics of the cohorts diagnosed before and after January 2005 and the likelihood of death within 3, 6, 9, or 12 months of diagnosis. The goal in analyzing characteristics is to establish that individuals diagnosed just before relative to just after the payment change are similar on observable dimensions, a key identifying assumption for our analysis. For the survival analysis, we separately obtained dates of death from administrative records through June 30, 2012 to minimize any right truncation bias. Because of the low rate of lung cancer survival, this follow-up period captures almost 90% of deaths in both cohorts..

Consistent with the plots, we model the effect of the January 2005 switch to ASP-based payments using week rather than month of diagnosis as follows:

$$\bar{Y_t} = \beta_0 + \beta_1 * post + \beta_2 * (t - t^*) + \beta_3 * (t - t^*) * post + \bar{X}_t\gamma + \epsilon_{ct}, \tag{5}$$

where $\bar{Y_t}$ is an outcome such as the mean rate of chemotherapy treatment for beneficiaries diagnosed in week t, post is an indicator equal to 1 after the payment change and 0 otherwise, $(t - t^*)$ is a linear function of time relative to the reform, t^*, that controls for smooth trends in outcomes around the time of the payment change and that we allow to differ on either side of the payment change. These regressions control for the share of patients by week that were male, the share by race/ethnicity, the median age and median age squared, the mean Deyo-Charlson score, and the

[18] In principle we have 12 months of post-reform data but the cohort counts show a sharp drop off in December, suggesting the data were incomplete.

share with metastatic disease. We computed Newey-West standard errors, which allow for heteroskedasticity and autocorrelation in the error of an unknown form up to a lag of 52 weeks (Newey and West (1987)). We allowed for this long a lag to account for correlation of annual events that affect treatment, such as the Christmas holidays. However, inference is not meaningfully affected by alternative lag structures. In sensitivity checks, we estimate regressions that include month-relative to January 2005 trends or that omit these trends but include calendar month fixed effects to control for cyclicality in treatment timing.

4 Results

4.1 Descriptive Statistics

Table 1 shows demographic characteristics for our analytic sample overall and by diagnosis before versus after January 2005. As mentioned above, we focus on the 11 months before and after the payment change – from February 2004 to November 2005. Median age at diagnosis was 74 years. Just over half the patients were male, 88% were white, and 9% were black. Patients had an average Deyo-Charlson comorbidity score of 1.08, meaning one non-cancer comorbidity in the year leading up to diagnosis, and almost 29% had metastatic disease documented in claims within 1 month of diagnosis.[19] While the Charlson score and the proportion male are statistically different for the pre and post-reform cohorts, neither difference (0.02 more comorbidities or 0.4% fewer males) is meaningful in magnitude. There were no other differences in observable demographic characteristics.

We formally test the stability of patient characteristics, since our research design assumes newly diagnosed patients are similar just before and after ASP implementation, by estimating variants of equation (1) with these characteristics as dependent variables. The coefficients and 95 percent confidence intervals for the month relative to reform fixed effects are shown in Figure 1a-e. With the possible exception of the comorbidity score, which is noisy, these characteristics (e.g., share metastatic,

[19]We use a 1-month period to define metastasis since our models capture monthly treatment. Given this relatively short period, this measure should yield conservative estimates of metastasis.

white, male and log age) are smooth across the reform (time 0). The estimated discontinuities in these variables on January 2005 are in Appendix Table 1. As expected, the estimates tend to be small and indistinguishable from zero. When the estimates are distinguishable from zero (specifically, log age and the share of patients of other/unknown race), they are modest in size. To the extent that the gap in age is real, it should bias us against finding improvements in survival. On balance we interpret the results in Figure 1 and Appendix Table 1, as evidence that a key identifying assumption of the empirical approach – the stability of the population of newly diagnosed lung cancer patients across the reform – is satisfied.

Table 1 and Figure 2 provide evidence that the reform had bite by showing payment changes for several drugs commonly used to treat lung cancer. Payment rates declined after the reform for carboplatin, paclitaxel, and etoposide. The changes were most striking for carboplatin and paclitaxel; reimbursement rates for a standard monthly dose declined from over $2,270 to $225 for paclitaxel and from $1,845 to $930 for carboplatin. As shown in Figure 2, the changes in quarterly carboplatin and paclitaxel reimbursements occurred around the time the new payment system took effect (time 0)). The large discrepancy between payment rates and acquisition costs for these drugs was identified as early as 1997. In 2001, the General Accounting Office (now the Government Accountability Office) reported widely available discounts of about 20% below AWP for both drugs (Government Accountability Office (2001)); a later analysis showed that 2004 paclitaxel reimbursements were six times higher than cost (Government Accountability Office (2004)). Payments were relatively flat for docetaxel, a high-priced drug at about $2,500 per standardized monthly dose, and gemcitabine at $1,300 per monthly dose, indicating the reform differentially affected drugs, depending on the pre-reform mark-up.

A complication in the payment change is that carboplatin went off patent in October 2004, the quarter before ASP implementation. Because of the lag in Average Sales Price determination – quarterly reimbursements are based on wholesale prices two quarters prior – carboplatin likely maintained a margin above the 6% mandated by the legislation for at least the first quarter of 2005.[20] To ensure that our results for

[20]Upon its patent expiration in October 2004, at least 5 generic manufacturers had secured FDA approval to enter the carboplatin market. The gradual decline in carboplatin's payment rates in

chemotherapy are not driven entirely by carboplatin, where the observed changes in reimbursement combine the effects of the patent expiration and the payment reform, we examine both chemotherapy use of regimens with and without carboplatin.

Table 2 provides summary statistics for our key outcomes of interest. These descriptive statistics indicate significant changes in the likelihood of chemotherapy treatment after the payment change. Whereas 16.6% of patients diagnosed in the 11 months prior to the reform received chemotherapy within 1 month of diagnosis, those diagnosed in the 11 months after were almost 2 percentage points (or 11% off the pre-reform base) more likely to receive chemotherapy within 1 month of diagnosis. Nearly three-quarters of this increase (1.5 percentage points) was among patients receiving regimens without carboplatin. In addition, this increase came almost entirely from treatment in physician offices rather than outpatient hospital clinics, which were not subject to a reimbursement change in 2005. The relative increase in chemotherapy treatment was similar within 3 months (a 3.2 percentage point increase off a pre-reform base of 26%) as was the composition of the changes across settings.

Conditional on treatment within 1 month of diagnosis, the percent of patients receiving carboplatin declined from almost 56 to 53, and the percent receiving paclitaxel declined from 30 to 26, consistent with the large payment decline for these drugs. Since carboplatin and paclitaxel are often given in combination and because carboplatin likely retained a margin above 6% for a few quarters, some of carboplatin's decline may be driven by the sharp reduction in paclitaxel's profitability. Trends in use of etoposide were comparatively flat while use of docetaxel, a relatively high-priced alternative to palitaxel, increased 1.3 percentage points off a base of just 8.3%, possibly reflecting the incentive from the uniform 6% margin on all drugs.[21]

Among cohorts diagnosed in the 11 months prior to January 2005, 33.7% died within 3, 46.7% within 6, 52.4% within 9, and 61.6% within 12 months of diagnosis.

Figure 2, which may capture switching to generics, suggests it maintained a margin above 6% for longer than the first quarter of 2005.

[21] Although the rise in prescription drug shortages began around 2005, the rise in oncology drug shortages occurred much later, around 2008. For details see http://healthaffairs.org/blog/2012/05/29/prescription-drug-shortages-reconsidering-the-role-of-medicare-payment-policies/ Paclitaxel and Carboplatin were not in short supply until 2009 and 2010, respectively. Thus, shortages are unlikely to explain the drug switching observed here.

For cohorts diagnosed in the 11 months after January 2005, the proportion dying was 1-2 percentage points (3%-4%) lower in each of these four periods.

4.2 Changes in Treatment

Figures 3a-3d show the regression-adjusted mean changes in the likelihood that patients received any chemotherapy, any chemotherapy regimen excluding carboplatin, any chemotherapy in a physician's office, and any chemotherapy in an outpatient clinic by month of diagnosis relative to January 2005. With the exception of the outpatient clinic setting (Fig 3d), which did not face a payment change until 2006, all the panels show a discrete increase in the likelihood of chemotherapy treatment for lung cancer patients diagnosed just after relative to just before January 2005 (time 0). The timing of the increase and that most of the increase came in physician offices, the only setting subject to the payment reform in 2005, suggest that the switch to the ASP system drove the change. The similarity when carboplatin is excluded confirms that the increase in treatment is not driven by the sustained margin on this drug.

Table 3 shows estimates of the discontinuous jump in chemotherapy treatment for those diagnosed just after January 2005. The estimates are based on the specification in (5) using 11 months (col 2) or 9 months (col 3) on either side of the payment change. The overall change in the likelihood of chemotherapy treatment rates within 1 month of diagnosis is about 1.5 percentage points. Off a base of 16.6% of those diagnosed prior to the reform who receive treatment within one month, this represents about a 9% increase in the likelihood of treatment in response to the reform. When we look at treatment that does not include carboplatin, the relative increase is much larger – almost 20% (1.2 to 1.4 percentage points off a base of 7.3%). This implies that the overall change in chemotherapy treatment is not driven solely by the interaction between the payment reform and the patent expiration of carboplatin. In addition, the increase is driven by treatment in the physician's office setting, which was the setting subject to the reform. As shown in Appendix Table 2, the results are largely insensitive to the use of relative month trends (cols 1 and 2) or calendar month fixed effects (cols 3 and 4) instead of relative-week trends. The main difference is that the effects are slightly larger when calendar month instead of relative-week or relative-

month trends are used, suggesting a 10% increase in treatment and an increase in treatment in the outpatient setting. However, the patterns in Figure 3 suggest that the specification with relative trends that are allowed to vary on either side of January 1, 2005 more appropriately captures treatment trendfs before and after the reform.

Figures 4a-4d and Table 4 consider chemotherapy service counts within thirty days of diagnosis. Changes in service counts, unlike the likelihood of treatment, combine both extensive and intensive margin changes. The trends in Figure 4 are analogous to those in Figure 3 but appear larger, at least outside of the outpatient clinic setting, where the change was essentially zero. As shown in Table 4, the mean number of chemotherapy services within 30 days was 1 prior to the reform; among those diagnosed after the reform, services increased almost 33%. Given the 9% increase in treatment on the extensive margin, these results suggest that chemotherapy services likely increased even among those who would have received treatment absent the reform. As on the extensive margin, the increase occurs in patients receiving treatments that do not contain carboplatin, implying that the interaction of the reform and carboplatin's patent expiration cannot fully explain our main treatment result. The increase in physician offices accounts for almost all (about 94% or 0.916 *0.347/0.339) of the change. Services increase somewhat in the outpatient clinic setting as well, about 1.3%, although this change is modest compared to the 34% increase in physician offices. As with the likelihood of any treatment, models with relative-month trends or calendar month fixed effects (Appendix Table 3) yield results that are similar, albeit larger without linear time trends.

As a more formal test of a structural break in treatment, we adopt an approach from the time series literature. Specifically, we employ a Quandt likelihood ratio (QLR) test (Quandt 1960) that tests for structural breaks at all possible dates between two time periods, τ_0 to τ_1. Operationally, the test is a modified set of sequential Chow tests based on equation (2), where a post-period is defined at each cutoff date τ between the two time periods, i.e. $post = 1(t \geq \tau^*)$ is an indicator equal to zero for all weeks before τ^* and one for all subsequent weeks; this indicator is also interacted with the linear trend. Under the null hypothesis of no break, the coefficients on $post_\tau$ and $post_\tau * (t - t^*)$ are jointly equal to zero. The QLR statistic, also known as the sub-Wald statistic, is just the maximal of the F-statistics testing the null of no break

at all dates between τ_0 to τ_1; it provides an estimate of the structural break date. To implement this test, we adopt the conventional choice for τ_0 to τ_1 of 0.15T and 0.85T or -34 weeks and + 33 weeks relative to the first week, i.e. week 0, in January 2005. In this way, the post period cutoff or break date, τ^*, is sequentially defined over the inner 70% of the 96 week sample. We take the critical values of the QLR statistic for the endpoints 0.15T and 0.85T and two restrictions tested from Stock and Watson (2003), which are adapted from Andrews (2003).[22]

Appendix Figure 1 plots the F-statistics testing for a break at -34 to +33 weeks relative to week 0. We show these tests for the following 30-day treatment series: any chemotherapy, chemotherapy that does not include carboplatin, chemotherapy in a physician's office, and chemotherapy in an outpatient hospital setting. Based on the QLR statistic, a structural break in treatment occurs two weeks prior to January 2005 for any chemotherapy treatment (App Fig 1a) and treatment in a physician's office (App Fig 1c). Excluding Carboplatin, the break occurs 1 week prior to ASP implementation (App Fig 1b). The pre-implementation break is not worrisome given that we are analyzing 30-day treatment and patients diagnosed at the end of 2004 likely began treatment after the new payment system took effect on January 1 rather than during the Christmas holidays. In all three cases, the maximal F-statistic exceeds the 1% critical value of 7.78, implying we can reject the null of no structural break in treatment. In contrast, we cannot reject the null at any point for chemotherapy treatment in the outpatient setting (Appendix Fig 1d).

The results above indicate that chemotherapy treatment within 30 days of diagnosis increased on both the extensive and intensive margins. As another check, we analyze the likelihood of treatment within 90 days of diagnosis. The 90 day period captures those diagnosed just before the reform but who were affected by the reform if they survived past implementation. Consistent with this, Appendix Figures 2a-2c show a gradual increase in 90-day treatment for those cohorts diagnosed within three to one months prior to the reform. The change flattens out for those cohorts diagnosed just after the reform, suggesting that the earlier increase captures partial exposure to the reform. As indicated in Appendix Table 4, once this gradual rise is

[22]The QLR critical values are larger than the critical value for a single F-test; they are effectively corrected for multiple-hypothesis testing induced by searching over many F-tests.

accounted for (via an indicator for diagnosis in the quarter prior to the reform), the increase in 90-day treatment for those fully exposed to the reform, i.e. diagnosed in 2005, was about 11% overall and almost 15% in the physician's office-setting. In a parallel fashion, Figure 3d shows that cohorts diagnosed in the last few months of 2005 experienced an increase in the likelihood of treatment in the outpatient setting, after the payment reform took effect in this setting. In other words, the treatment response to the payment change in the physician's setting is not unique.

We next analyze changes in the types of drugs used. Figure 5a-5e show regression-adjusted changes in the probability that chemotherapy-treated patients received carboplatin, paclitaxel, docetaxel, etoposide and gemcitabine, the most commonly prescribed agents in our sample. Patients treated within 30 days of diagnosis were less likely to receive a mix of agents that included carboplatin or paclitaxel. The timing of the decline in use of these agents preceded the introduction of the average sales price payment system by about 3 months, which may reflect the quarterly purchasing common among oncologists.[23] Given the publicity the reform received among oncologists, physicians likely understood that the reduction in payment rates for these drugs would be large and were reducing their inventory of these agents in advance. Failure to do so could have meant a considerable loss of income. In contrast to the pattern for paclitaxel, the probability of receiving docetaxel, an expensive agent implicitly favored by the 6% margin on all drugs, increased modestly for patients receiving chemotherapy treatment. The increase preceded the reform by about a month, further evidence that physicians were rearranging their stock of agents in anticipation of the ASP system. The likelihood of receiving etoposide or gemcitabine and other less commonly used agents did not change systematically.

Table 5 quantifies the changes illustrated in Figure 5 but first considers the unconditional change in use within 30 days. Whereas carboplatin use did not change unconditionally, use of paclitaxel, the drug that lost the most margin, declined by about 10% and docetaxel and etoposide increased by about 30% and 9%, respectively. These changes combine the change in chemotherapy treatment and the choice of drugs. Conditional on any treatment, physicians were less likely to give patients carboplatin

[23]We thank the community oncologist from Phoenix, AZ who first relayed this fact to us.

or paclitaxel after ASP implementation. The probability that chemotherapy-treated patients received carboplatin declined 3 to 4 percentage points. The decline in the probability of receiving paclitaxel was more dramatic – about 6 to 7 percentage points or over 20%. Some physicians switched patients to docetaxel, an expensive agent favored by the 6% uniform mark-up. The docetaxel increase was about 2.6 percentage points or over 30% off a base of 8.3%. Accounting for pre-reform switching increases the magnitude of these changes (see Appendix Table 5). The change for etoposide is small, indistinguishable from zero and changes sign as the study window narrows. The estimate for gemcitabine is larger in relative terms (0.7 percentage points off a base of 10%), but based on Fig 5e may not reflect a policy-related change.

4.3 Changes in Survival

The results above demonstrate that chemotherapy treatment increased on both the intensive and likely extensive margin for lung cancer patients diagnosed just after relative to just before the reform. The mix of drugs administered also changed, with drugs that lost the most margin used less and expensive agents favored by the 6% average margin used slightly more. Although cancer treatment is constantly evolving, no other change in management was so suddenly implemented that could account for changes of the magnitude we observed. Evidence for adjuvant chemotherapy was emerging in the mid-2000s but the timing of major presentations and publications on this subject are unlikely to have caused a sudden change in practice in January 2005 (Le Chevalier (2003); Arriagada R. et al. (2004); Winton T. et al. (2004); Winton T. et al. (2005); Struass G. et al. (2004); Struass G. et al. (2006); Struass G. et al. (2008); Douillard J. et al. (2005); Douillard J. et al. (2006)). Furthermore, evidence on uptake of adjuvant chemotherapy suggests utilization was stable across the policy change (Booth C. M. et al. (2010)), and our own data (not shown) show no increase in the use of cisplatin-based regimens, as would be expected if the diffusion of adjuvant chemotherapy were to explain our findings.

The treatment changes shown here beg the question of what happened to patient welfare. We have one important measure of welfare – survival – that we exploit to address this issue. A limitation is that we cannot measure changes in quality

of life, which could deteriorate in response to harsh chemotherapy treatment (e.g., see Ballatori et al. (2007)). Despite the common view that chemotherapy decreases quality of life, however, a recent randomized trial in patients with advanced lung cancer found that, compared to placebo, chemotherapy not only increased survival but also improved some measures of quality of life (e.g., the worsening of pain and coughing up of blood) while leaving others unchanged (Belani C. P. et al. (2012)).

To begin, Appendix Figure 3 plots the unadjusted Kaplan-Meier survival curves for patients diagnosed in the 11 months before and after the reform. The median survival based on the Kaplan-Meier estimator was 261 and 284 days for those diagnosed in the pre and post-reform periods, respectively, a difference of about 3 weeks. To assess whether this difference is related to the reform, Figures 6a-6d plot by month of diagnosis relative to the January 2005 reform, the regression-adjusted mean changes in the proportion of patients dying within 3, 6, 9 and 12 months of diagnosis. These figures show a discrete decline in the likelihood of dying within each interval for the post-reform cohorts. The decline, at least for death within 3 and 9 months, corresponds closely to the timing of ASP implementation and the previously demonstrated increase in chemotherapy utilization.

Table 6 quantifies the mortality changes in Figure 6a-6d. Among patients diagnosed just after relative to just before the reform, the likelihood of death within 3 months of diagnosis decreased about 1 percentage point or about 4% relative to the base of 33.9% dying in this interval. The results are virtually identical using patients diagnosed 9 months before to 9 months after the payment change (col 3) or using alternative specifications (see Appendix Table 6). At 6 months from diagnosis, the reduction in the likelihood of death is about 1.6 percentage points or 3.4% relative to the base rate of 46.9%. At 9 months the effect is 1.1 percentage points, a relative decrease in mortality of about 2.1%. At 1 year, the survival effects seem to disappear. While the reductions in the likelihood of death at 3 to 9 months post-diagnosis are seemingly small, these estimates average survival across all patients, whereas the reform increased chemotherapy in less than 2 out of every 100 patients. The reform also changed the intensity of treatment and the drugs used, making it difficult to isolate the specific source of the survival change.

We take several approaches to testing whether the survival changes were caused

by the reform rather than capturing patterns unaccounted for by the week trends in the two periods. First, we take advantage of substantial geographic variation in the response to the payment change. Jacobson et al. (2011) shows that after the reform the likelihood of chemotherapy treatment was virtually unchanged in some states (e.g., CA and MO), increased more than twice the national average in others (e.g., MN, CT) and even declined in a few states (e.g., OK, ID). If reform-related changes in chemotherapy improved survival, these effects should be concentrated in the most responsive states. To this end, Figures 7a-7d show the likelihood of death within 3-12 months of diagnosis for patients in states with above median 30-day treatment response; Figures 7e-7h show analogous plots for states with below median response. As expected, declines in the likelihood of death at 3, 6, and 9 months are greater for the above-median response states than for the whole sample. The below-median response states show no clear break in survival for cohorts diagnosed just after relative to just before the reform. Table 7 confirms the visual patterns: while the likelihood of death declined by 3%-6% within 3-6 months of diagnosis and 1.5%-3% within 9 months of diagnosis in the most responsive states survival declined, if anything, in the least responsive states. More plausibly, based on the plots and the estimated changes within 6-9 months of diagnosis, survival in the last responsive states was unaffected by the reform. These results support the view that the increase in chemotherapy generated by the reform improved survival.

As a second approach, we stratify states based on whether they had above or below median pre-reform rates of paclitaxel use in chemotherapy-treated patients. Heavy reliance on paclitaxel pre-reform meant that the payment change had larger bite, i.e. doctors in these areas had greater exposure to the payment cuts. As shown in Appendix Figures 4a-4c, 30-day chemotherapy treatment changes are quite sharp in states with high pre-reform reliance on paclitaxel. Treatment also increased in states with lower pre-reform paclitaxel use (see Appendix Figures 4e-4g) but the plots are considerably noisier and the estimates in Appendix Table 7 are more sensitive to the study window. Moving from the 11 to 9-month study window, the estimated treatment changes increase slightly in above-median states but fall by as much as 65% in below-median states. Based on 9-months on either side of the reform, 30-day chemotherapy treatment rates increased by 2.4 percentage points or 14% in states with

above-median pre-reform use of paclitaxel but only 3% in states with below median use. These treatment differences, which are statistically significant using the 9-month study window, translate into survival differences, as shown in Appendix Figures 5a-5h and the second panel of Appendix Table 7. The decline in the likelihood of dying within 3-9 months is sharper in the above versus below median pre-reform paclitaxel use states. While the likelihood of death declined by 4%-6% within 3-9 months in the above-median states, the estimated changes for the below-median states are small, often positive and, with the exception of one estimate at 1-year out, never statistically distinguishable from zero. Across states with high and low pre-reform rates of paclitaxel use, the differences in survival are statistically distinguishable at 3 months when using the short study window and at 6 and 9 months for both the long and short study windows.[24]

As a final approach to assessing the credibility of the survival effects, we split the sample by above and below median age patients. Since older patients are more likely to be undertreated (Davidoff et al. (2010); Booth C. M. et al. (2010)), they should benefit most from increased treatment. Appendix Figures 6a-6d and 6e-6h show changes in the likelihood of chemotherapy use for above and below-median age patients, respectively. Appendix Figures 7a-7d and 7e-7h show changes in the likelihood of death, for the same groups. Although treatment increases sharply for both groups (in all but the outpatient setting), a sharp reform-related decline in survival is clearest for the above-median age patients. Table 8 and Appendix Table 8 show the estimated changes in treatment and survival. Above-median age patients, who are almost half as likely to receive chemotherapy within 30 days, experience about an 8% increase in the likelihood of treatment; the increase is about 11% for the younger group. In contrast, the survival effects are substantially larger for the older group. The likelihood of death within 3 months declines by about 0.6 percentage points for the younger group but 2.3 percentage points for the older group. At 6 and 9 months of diagnosis, the declines are 2.6-3.2 percentage points in the older group and 1.2-2 percentage points in the younger group. Relative to baseline survival, these changes represent a decrease in the likelihood of death within 3-9 months of 5-6%

[24]Using the long study window, the 6-month difference is only distinguishable at the 10% level. All other 6-9 month differences are significant well below the 5% level.

for the older group compared to 2-4% for the younger group. The differences in treatment and survival at 3-9 months across above and below median age patients are statistically distinguishable.[25]

4.4 Market Concentration Interactions

To test whether the changes in treatment and survival are related to market concentration, we compare the response of low concentration (i.e. more competitive) and high concentration (i.e. less competitive) markets in response to the policy change. Our primary measure of market concentration is a county level Herfindahl-Hirshman Index (HHI) for chemotherapy providers. Specifically we calculate the HHI as the sum of each provider's share of a county's chemotherapy administrations *pre-reform*.[26] Following the US DOJ Horizontal Merger Guidelines (2010), we define a county as high concentration if the HHI is greater than 0.25. As an alternative measure, we calculate the pre-reform ratio of patients to providers in a county, and define a country as highly concentrated if a county's ratio is above the median value for all counties.

We show the change in chemotherapy treatment for patients in high concentration counties in Figures 8a-8d and for medium/low concentration counties in Figures 8e-8h. On the extensive margin, we see no compelling evidence of a change in treatment probabilities in high concentration markets but a sharp increase right after the reform in less concentrated counties. When we consider service counts, a mix of intensive and extensive margin changes, we see a small increase in treatment in more concentrated markets (Appendix Figures 8a -8d) and again a sharper and larger increase in less concentrated markets. These visual patterns are confirmed in Table 9 and Appendix Table 9, where estimates of the change in chemotherapy treatment on the extensive margin are small and generally insignificant in high concentration counties but imply about a 12% increase in 30-day treatment probabilities in less concentrated counties. Chemotherapy service counts increase by about 20% in more concentrated markets

[25] Using the 9-month window, the chemotherapy difference is only distinguishable at the 10% level.

[26] Since we have any billing-provider, irrespective of whether she is an oncologist, we restrict to those who billed for chemotherapy for any patient in the unrestricted sample of over 800,000 beneficiaries. Because we have only a patient's county, providers can contribute to the HHI of multiple counties.

– indicating some change in treatment patterns in response to the reform – but by almost 40% in less concentrated markets. The differences in treatment rates and service counts are in all cases statistically different across market types.

In contrast to the more modest treatment changes in highly concentrated markets, Figures 9a-9h show that the likelihood of death within 3-9 months declined in both market types. The last panel of Table 9 shows that survival improved more in both absolute and relative terms in more concentrated markets. The likelihood of death within 6 months declines by about 2.2 percentage points or 4.6% in highly concentrated counties but 1.4 or 3% in less concentrated counties. This survival difference across markets is statistically significant, although only using the long study window. At 9 months the difference is starker, with a 1.9 percentage point or 3.5% decline in highly concentrated markets and a 0.007 percentage point or 1.3% decline in competitive markets. This difference is statistically significant for both study windows.

These patterns are broadly confirmed using a less formal definition of concentration – above vs. below median patients per providers. The likelihood of chemotherapy treatment increases modestly in counties with above median patients per providers (Appendix Figures 9a-9d) but more starkly in more competitive counties with below median patients per providers (Appendix Figures 9e-9h). As shown in Appendix Table 10, the extensive margin (any treatment) increases are about 1.3 percentage points or 7.5% in more concentrated (i.e., above median patients per provider) counties and 1.7% or 11% in less concentrated (i.e., below median patients per provider) counties, although these differences are not statistically distinguisable. The same basic pattern is found for service counts (Appendix Figures 10a-10h). The comparison for survival is flipped. The visual patterns show a break in survival in more concentrated counties (Appendix Figures 11a-11d) but seem to reflect cyclical changes in less concentrated counties (Appendix Figures 11e-11h). The last panel of Appendix Table 10 confirms these assessments, though the estimates are somewhat unstable across study windows. Using the narrow 9-month window, which more credibly isolates the impact of the reform, the likelihood of death within 6 months of diagnosis declines by 2.2 percentage points or 4.6% in more concentrated counties but by an insignificant 0.3 percentage points or 0.6% in less concentrated counties. At 9 months from diagnosis, the likelihood of death declines by 5% in more concentrated markets but is positive

(but insignificant) in less concentrated markets. These differences in survival at 3, 6 and 9 months are statistically distinguishable across market types.

5 Conclusion

In this paper, we find that among U.S. oncologists, physician agency is an important determinant of both the supply of heath care and patient outcomes. We find that a decrease in profit margins led to an increases in care, as measured by chemotherapy rates on both the intensive and extensive margins. The decrease in profit margins for chemotherapy drugs also led physicians to change the mix of drugs they use, shifting away from drugs that were more profitable before the reform to those that were relatively more profitable post-reform.

More surprisingly, given the widespread assumption that the key issue with physician agency is the inducement of unwarranted demand, we find that the increases in chemotherapy care improved survival. Cross-sectionally, the treatment and survival increases were larger in areas where the pre-reform mix of chemotherapy drugs meant a larger average decrease in margins from the reform. These results indicate that, in the presence of significant physician agency effects, an over-generous payment scheme can lead to physician rationing - i.e. the underprovision of care - and that better aligning payments to costs can lead to better outcomes at lower total cost. While we discuss under-provision in terms of rationing, the results are also consistent with doctors mistakenly believing they were providing the optimal amount of care prior to the reform and thus intentionally providing excess care post-reform. This alternative scenario is consistent with our model, but, if true, suggests an additional policy lever. Specifically efforts to promote evidence-based medicine might have improved outcomes even in the absence of payment reform, although without the cost-savings generated by the reform.

We also investigated the interaction between physician agency and the competitive environment. Consistent with our model, we found that increases in care were larger in more competitive (i.e. less concentrated) markets, while increases in survival rates were higher in less competitive (i.e. more concentrated) markets. This result suggest that while competition among providers can lead to less rationing of care, it also

makes them more responsive to financial incentives.

References

Akscin, J. and T. R. Barr (2007). Key practice indicators in office-based oncology practices: 2007 report on 2006 data. *Journal of Oncology Practice 3*(4), 200–203.

Andrews, D. W. (2003). Tests for parameter instability and structural change with unknown change point: A corrigendum. *Econometrica 71*(1), 395–397.

Arriagada R. et al. (2004). Cisplatin-based adjuvant chemotherapy in patients with completely resected non-small-cell lung cancer. *New England Journal of Medicine*, 351–360.

Ballatori, E., F. Roila, B. Ruggeri, M. Betti, S. sarti, G. Soru, G. Cruciani, M. D. Maio, B. Andrea, and R. Deuson (2007). The impact of chemotherapy-induced nausea and vomiting on health-related quality of life. *Supportive Care in Cancer 15*(2), 179–185.

Barlas, S. (2011). Hospital outpatient pharmacies win higher medicare payments for 2011. *Pharmacy and Therapeutics 36*(1), 11–21.

Belani C. P. et al. (2012). Quality of life in patients with advanced non-small-cell lung cancer given maintenance treatment with pemetrexed versus placebo (h3e-mc-jmen): Results from a randomised, double-blind, phase 3 study. *The Lancet Oncology 13*(3), 292–299.

Berndt, E. and J. P. Newhouse (2012). Pricing and reimbursement in u.s. pharmaceutical markets. In P. M. Danzon and S. Nicholson (Eds.), *The Oxford Handbook of the Economics of the Biopharmaceutical Industry*, pp. 201–265. Oxford UP.

Booth C. M. et al. (2010). Adoption of adjuvant chemotherapy for non-small-cell lung cancer: a population-based outcomes study. *Journal of Clinical Oncology 28*(21), 3472–3478.

Butcher, L. (2008, April). Insurers and oncologists forge a better cancer drug policy. *Managed Care Magazine*.

Chandra, A., D. M. Cutler, and Z. Song (2012). Who ordered that? the economics of treatment choices in medical care. In P. P. Barros, T. G. McGuire, and M. V. Pauly (Eds.), *Handbook of Health Economics*. Amsterdam: North-Holland.

Charlson, M. E., P. Pompei, K. L. Ales, and C. R. MacKenzie (1987). A new method of classifying prognostic comorbidity in longitudinal studies: Development and validation. *Journal of Chronic Disease 40*(5), 373–383.

Christensen, S. (1992). Volume responses to exogenous changes in medicare's payment policies. *Health Services Research 27*(1), 65–79.

Clemens, J. and J. Gottlieb (2012). Do physicians' financial incentives affect medical treatment and patient health? SIEPR Discussion Paper 11-017.

Colla, C. H., N. E. Morden, J. S. Skinner, J. R. Hoverman, and E. Meara (2012). Impact of payment reform on chemotherapy at the end of life. *Journal of Oncology Practice 8*(3s), e6s–e13s.

Community Oncology Alliance (2006). Community oncology alliance gathers in washington to address crisis in access to cancer care. *Medical News Today*.

Davidoff, A. J., M. Tang, B. Seal, and M. J. Edelman (2010). Chemotherapy and survival benefit in elderly patients with advanced non-small-cell lung cancer. *Journal of Clinical Oncology 28*, 2191–2197.

Deyo, R. A., D. C. Cherkin, and M. A. Ciol (1992). Adapting a clinical comorbidity index for use with icd-9 cm administrative databases. *Journal of Clinical Epidemiology 45*(6), 613–619.

Douillard J. et al. (2005). Anita: Phase iii adjuvant vinorelbine (n) and cisplatin (p) versus observation (obs) in completely resected (stage i-iii) non-small-cell lung cancer (nsclc) patients (pts): Final results after 70-month median follow-up. *Journal of Clinical Oncology 23(16S June 1 Supplement)*(Abstr. 7013).

Douillard J. et al. (2006). Adjuvant vinorelbine plus cisplatin versus observation in patients with completely resected stage ib-iiia non-small-cell lung cancer (anita): a randomised controlled trial. *Lancet Oncology 7*(9), 719–727.

Earle, C. C., A. B. Nattinger, and A. L. Potosky (2002). Identifying cancer relapse using seer-medicare data. *Medical Care 40*(8 Supplemental), IV: 75–81.

Encaoua, D. and A. Jacquemin (1980). Degree of monopoly, indices of concentration and threat of entry. *Internation Economic Review 21*(1), 87–105.

Escarce, J. (1993). Medicare patients' use of overpriced procedures before and after the omnibus budget reconciliation act of 1987. *American Journal of Public Health 83*(3), 346–355.

Friedman J. Y. et al. (2007). The medicare modernization act and reimbursement for outpatient chemotherapy: do patients perceive changes in access to care? *Cancer 110*(10), 2304–2312.

Government Accountability Office (2001). Medicare physician fee schedule: Practice payments to oncologists indicate need for overall refinements. *GAO-02-53*.

Government Accountability Office (2004). Medicare chemotherapy payments: new drug and administration fees are closer to providers' costs. *GAO-05-142R*.

Grant, D. (2009). Physician financial incentives and cesarean delivery: New conclusions from the healthcare cost and utilization project. *Journal of Health Economics 28*(1), 244–250.

Gruber, J., J. Kim, and D. Mayzlin (1999). Physician fees and procedure intensity: the case of cesarean delivery. *Journal of Health Economics 18*(4), 473–490.

Hurley, J. and R. Labelle (1995). Relative fees and the utilization of physician's services in canada. *Health Economics 4*, 419–438.

Hurley, J., R. Labelle, and T. Rice (1990). The relationship between physician fees and the utilization of medical services in ontario. *Advances in Health Economics and Health Services Research 11*, 49–78.

Jacobson, M., C. C. Earle, and J. P. Newhouse (2010). How medicare's payment cuts for cancer chemotherapy drugs changed patterns of treatment. *Health Affairs 29*, 1391–1399.

Jacobson, M., C. C. Earle, and J. P. Newhouse (2011). Geographic variation in physicians' response to a reimbursement change. *New England Journal of Medicine 365*, 2049–2052.

Kaiser Family Foundation (2010). Kaiser family foundation medicare chartbook.

Le Chevalier, T. (2003). Results of the randomized international adjuvant lung cancer trial (ialt): cisplatin-based chemotherapy (ct) vs no ct in 1867 patients (pts) with resected non-small cell lung cancer (nsclc). *Proceedings of the American Society of Clinical Oncology 22*(Abstr 6).

McGuire, T. G. (2000). Physician agency. In A. J. Culyer and J. P. Newhouse (Eds.), *Handbook of Health Economics*, pp. 462–217. North-Holland.

McGuire, T. G. and M. V. Pauly (1991). Physician response to fee changes with multiple payers. *Journal of Health Economics 10*(4), 385–410.

MedPAC (2006). Report to the congress: effects of medicare payment changes on oncology services.

Mullen, P. (2007, June). The arrival of average sales price. *Biotechnology Healthcare*, 48–53.

Newey, W. and K. West (1987). A simple, positive semi-definite, heteroscedastic and autocorrelation consistent covariance matrix. *Econometrica 55*, 703–708.

Nguyen, N. X. and F. W. Derrick (1997). Physician behavioral response to a medicare price reduction. *Health Services Research 3*, 283–298.

Office of Inspector General (1997a, December). Department of health and human services. excessive medicare payment for prescription drugs, no. oei-03-97-00290.

Office of Inspector General (1997b, December). Department of health and human services. excessive medicare payment for prescription drugs, no. oei-03-97-00290.

Ramsey, S. D., J. F. Scoggins, D. K. Blough, C. L. McDermott, and C. M. Reyes (2009). Sensitivity of administrative claims to identify incident cases of lung cancer: a comparison of 3 health plans. *Journal of Managed Care Pharmacy 15*(8), 659–668.

Rice, T. (1983). The impact of changing medicare reimbursement rates on physician-induced demand. *Medical Care 21*, 803–815.

Rice, T. (1984). Physician-induced demand for medical care: New evidence from the medicare program. *Advances in Health Economics and Health Services Research 5*, 129–160.

Rice, T. and R. Labelle (1989). Do physicians induce demand for medical services? *Journal of Health Politics Policy Law 14*, 587–600.

Shainian, V. B., Y. Kuo, and S. Gilbert (2010). Reimbursement policy and androgen-deprivation therapy for prostate cancer. *New England Journal of Medicine 363*, 1822–1832.

Shea, A. M., L. H. Curtis, B. G. Hammill, L. D. DiMartino, A. P. Abernethy, and K. A. Schulman (2008). Association between the medicare modernization act of 2003 and patient wait times and travel distance for chemotherapy. *JAMA 300*(2), 189–196.

Skinner, J. (2012). Causes and consequences of regional variations in health care. In P. P. Barros, T. G. McGuire, and M. V. Pauly (Eds.), *Handbook of Health Economics*, Volume 2. Amsterdam: North-Holland.

Stock, J. H. and M. W. Watson (2003). *Introduction to Econometrics*. Pearson Education, Inc.

Struass G. et al. (2004, July). Randomized clinical trial of adjuvant chemotherapy with paclitaxel and carboplatin following resection in stage ib non-small cell lung cancer (nsclc): Report of cancer and leukemia group b (calgb) protocol 9633. *Journal of Clinical Oncology 22*(14S).

Struass G. et al. (2006, June). Adjuvant chemotherapy in stage ib non-small cell lung cancer (nsclc): Update of cancer and leukemia group b (calgb) protocol 9633. *Journal of Clinical Oncology 24*(18S).

Struass G. et al. (2008). Adjuvant paclitaxel plus carboplatin compared with observation in stage ib non-small-cell lung cancer: Calgb 9633. *Journal of Clinical Oncology 26*(31), 5043–5051.

Vera-Llonch, M., D. Weycker, and A. Glass (2011). Healthcare cost in patients with metastatic lung cancer receiving chemotherapy. *BMC Health Services Research 11*, 305.

Warren, J. L., E. Feuer, A. L. Potosky, G. F. riley, and C. F. Lynch (1999). Use of medicare hospital and physician data to assess breast cancer incidence. *Medical Care 37*(5), 445–456.

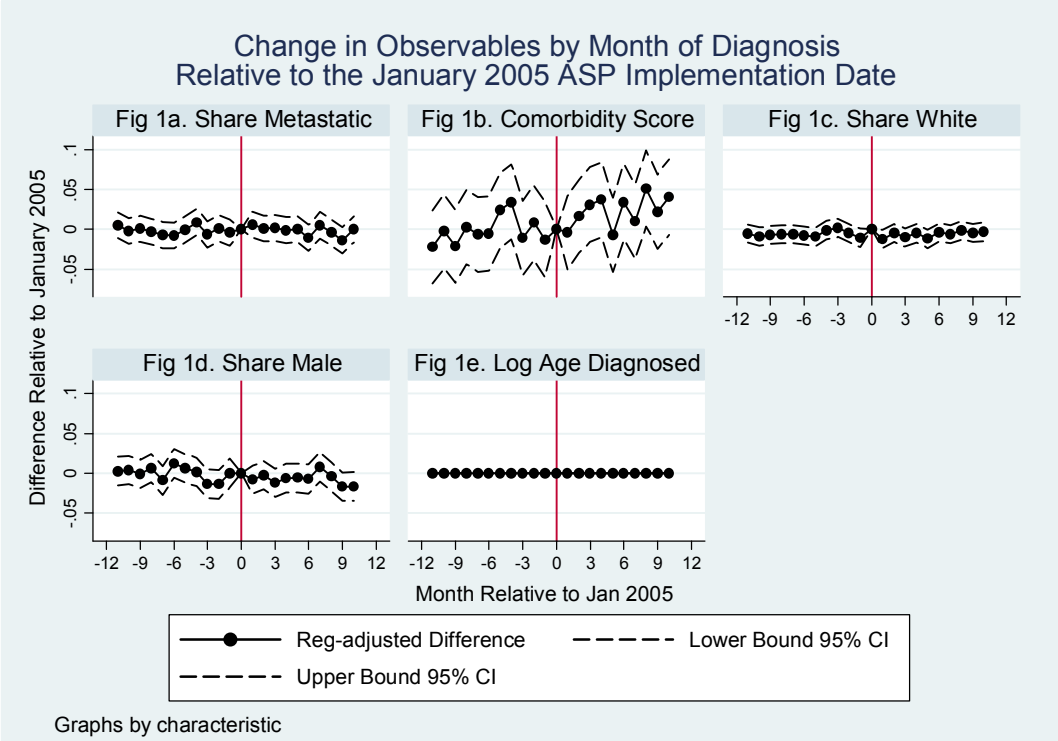
Warren, J. L., L. C. Harlan, A. Fahey, B. A. Virnig, J. L. Freeman, C. N. Klabunde, G. S. Cooper, and K. B. Knoft (2002). Utility of the seer-medicare data to identify chemotherapy use. *Medical Care 40 Supplemental*, IV:55–61.

Weight, C. J., E. J. Klein, and J. S. Jones (2007). Androgen deprivation falls as orchiectomy rates rise after changes in reimbursement in the u.s. medicare population. *Cancer 110*, 2304–2312.

Winton T. et al. (2004, July). A prospective randomised trial of adjuvant vinorelbine (vin) and cisplatin (cis) in completely resected stage 1b and ii non small cell lung cancer (nsclc) intergroup jbr.10. *Journal of Clinical Oncology 22*(14S).

Winton T. et al. (2005). Vinorelbine plus cisplatin vs. observation in resected non-small-cell lung cancer. *New England Journal of Medicine 352*(25), 2589–2597.

Yip, W. C. (1998). Physician response to medicare fee reductions: Changes in the volume of coronary artery bypass graft (cabg) surgeries in the medicare and private sectors. *Journal of Health Economics 17*, 675–699.


Change in Observables by Month of Diagnosis
Relative to the January 2005 ASP Implementation Date
Fig 1a. Share Metastatic
Fig 1b. Comorbidity Score
Fig 1c. Share White
Fig 1d. Share Male
Fig 1e. Log Age Diagnosed
Difference Relative to January 2005
-.05
0
.05
.1
-12 -9 -6 -3 0 3 6 9 12
Month Relative to Jan 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by characteristic

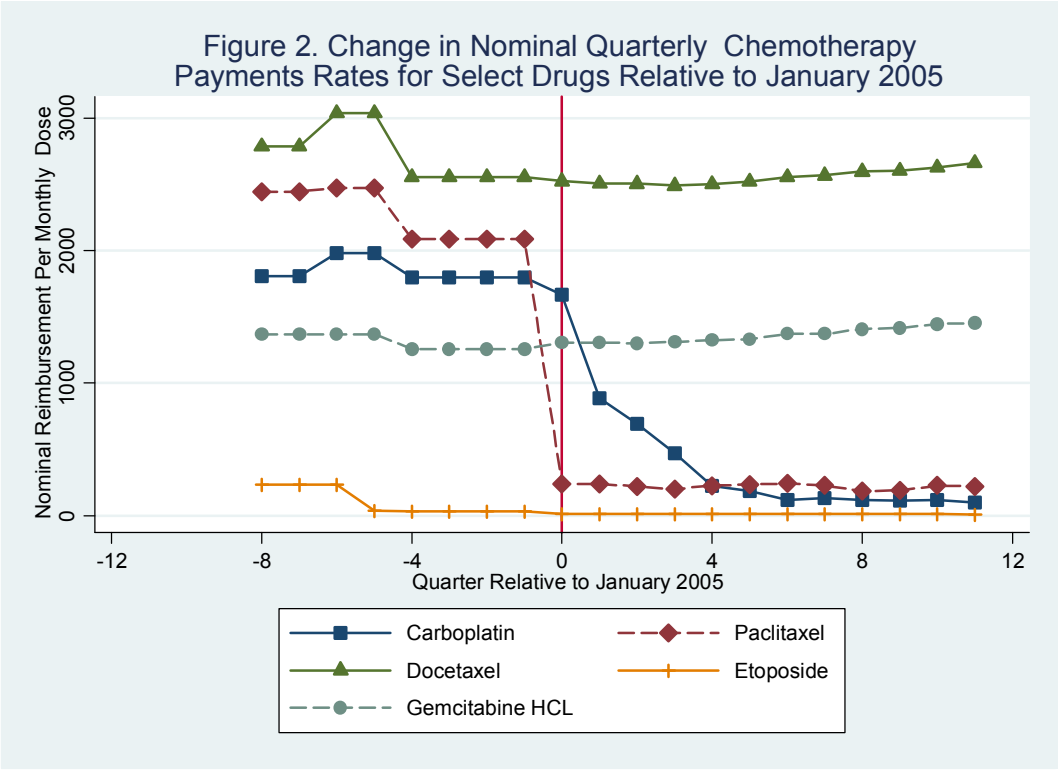
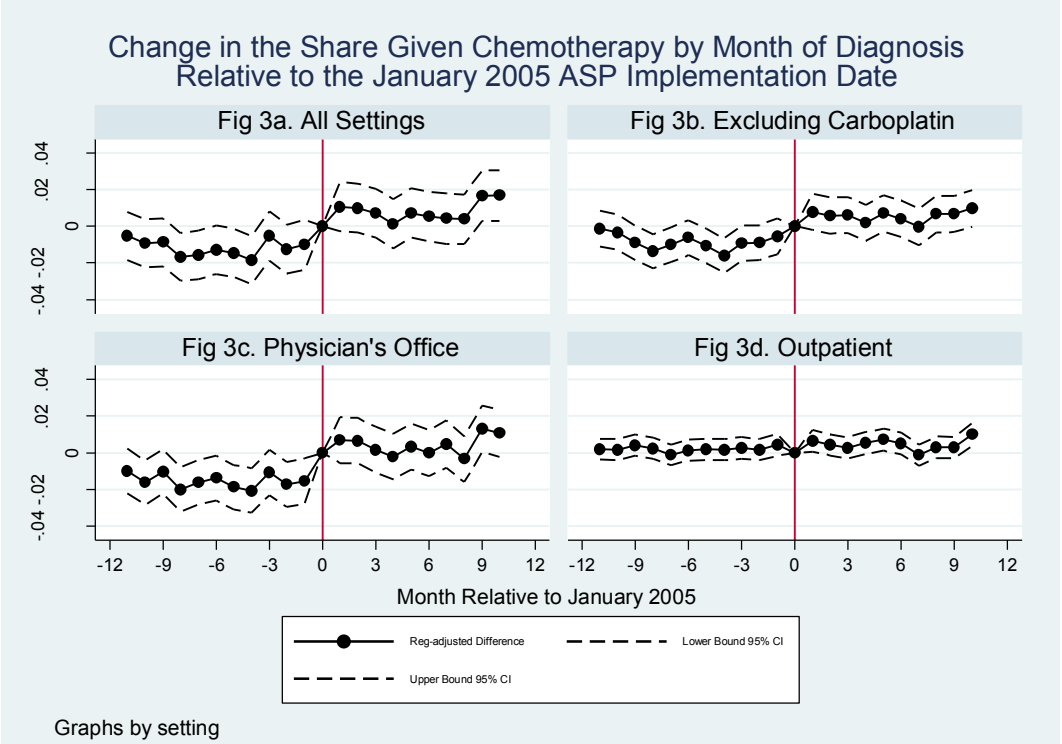
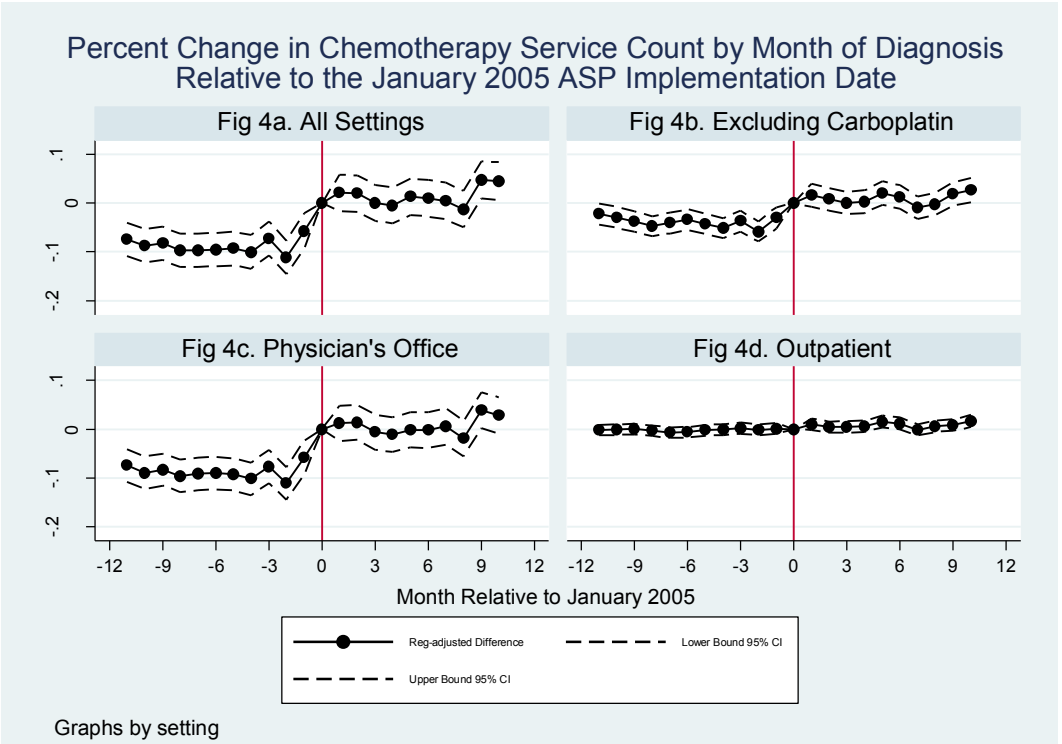
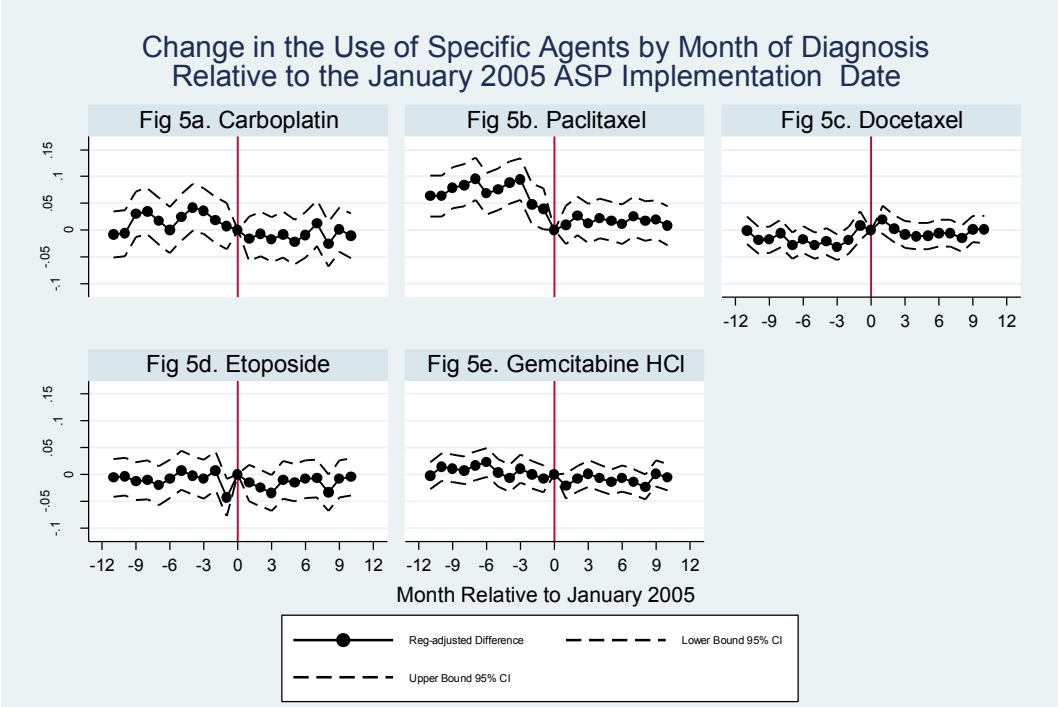
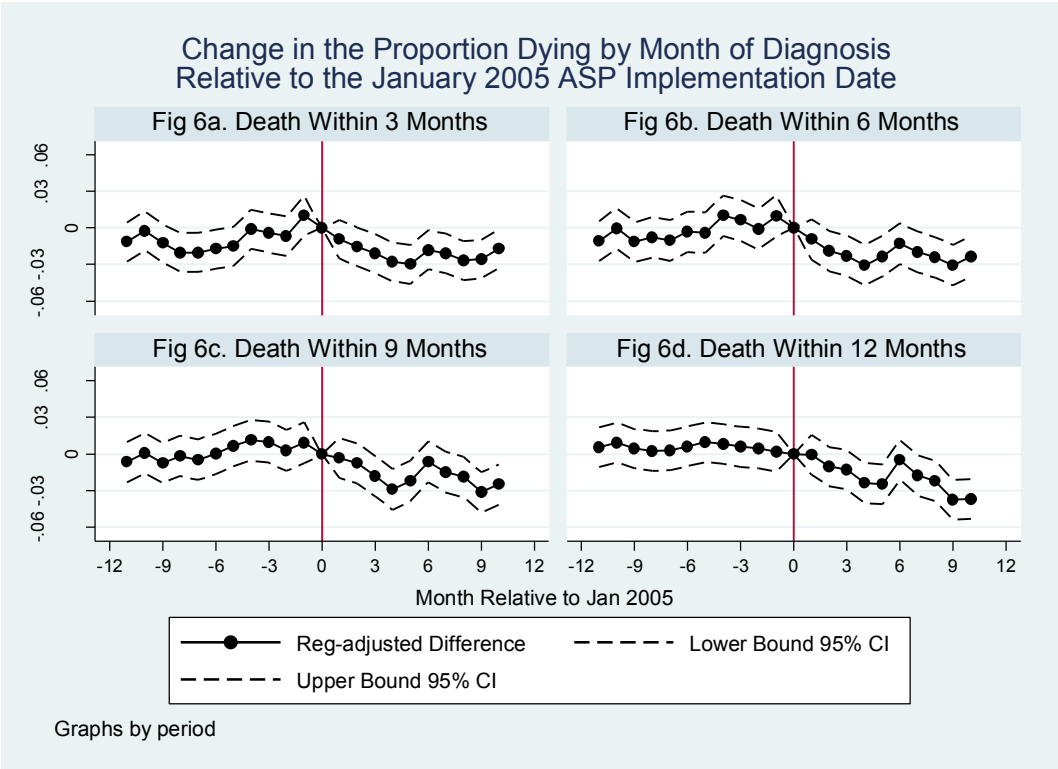


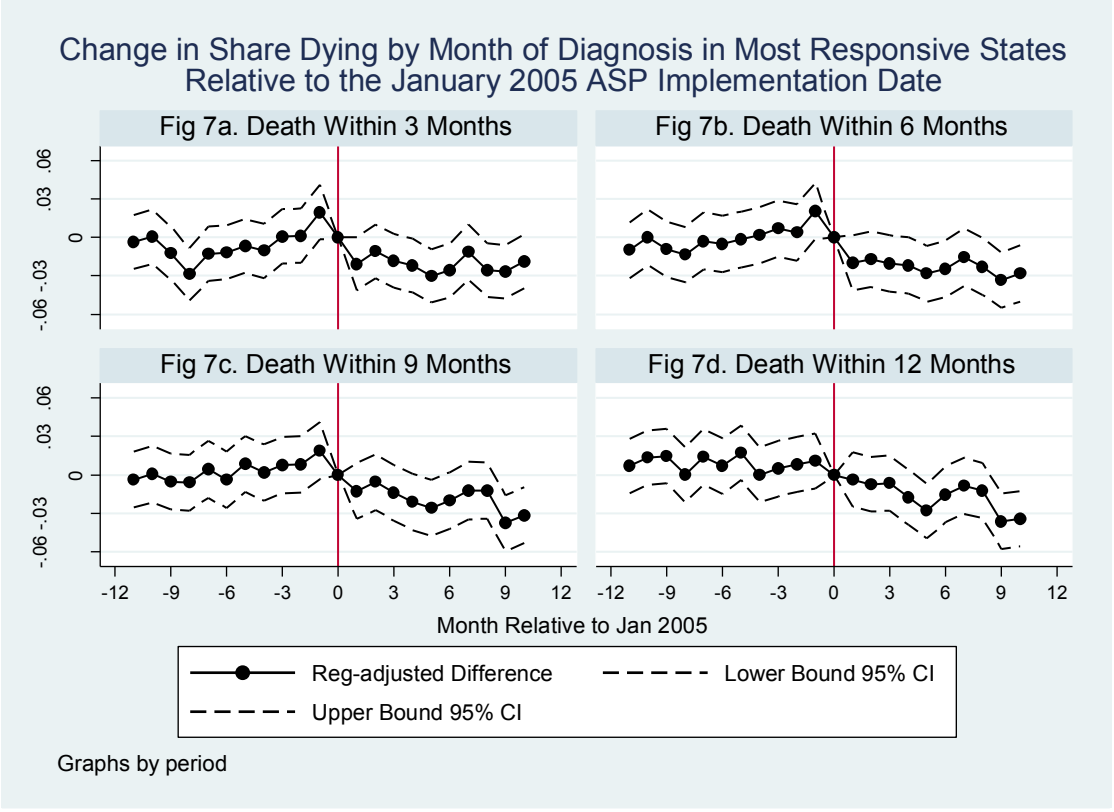
Figure 2. Change in Nominal Quarterly Chemotherapy
Payments Rates for Select Drugs Relative to January 2005
Nominal Reimbursement Per Monthly Dose
0
1000
2000
3000
-12
-8
-4
0
4
8
12
Quarter Relative to January 2005
Carboplatin
Paclitaxel
Docetaxel
Etoposide
Gemcitabine HCL

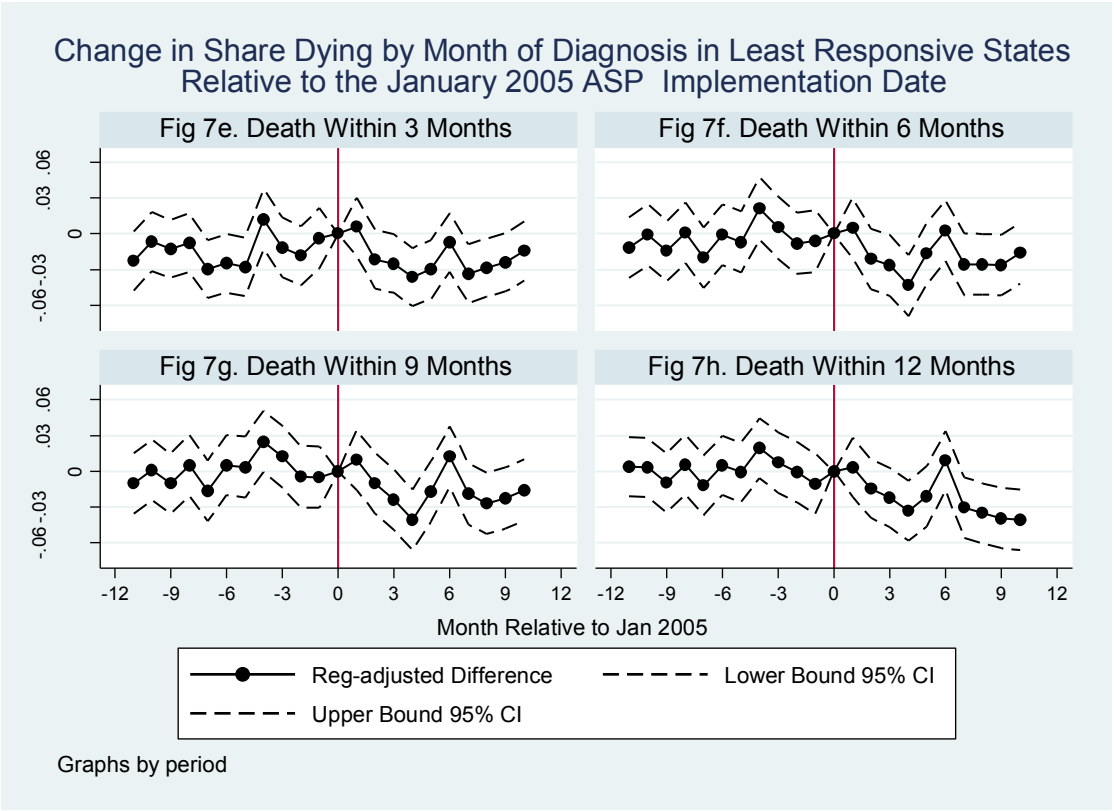

Change in the Share Given Chemotherapy by Month of Diagnosis
Relative to the January 2005 ASP Implementation Date
Fig 3a. All Settings
Fig 3b. Excluding Carboplatin
Fig 3c. Physician's Office
Fig 3d. Outpatient
-.04 -.02 0 .02 .04
-12 -9 -6 -3 0 3 6 9 12
Month Relative to January 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by setting

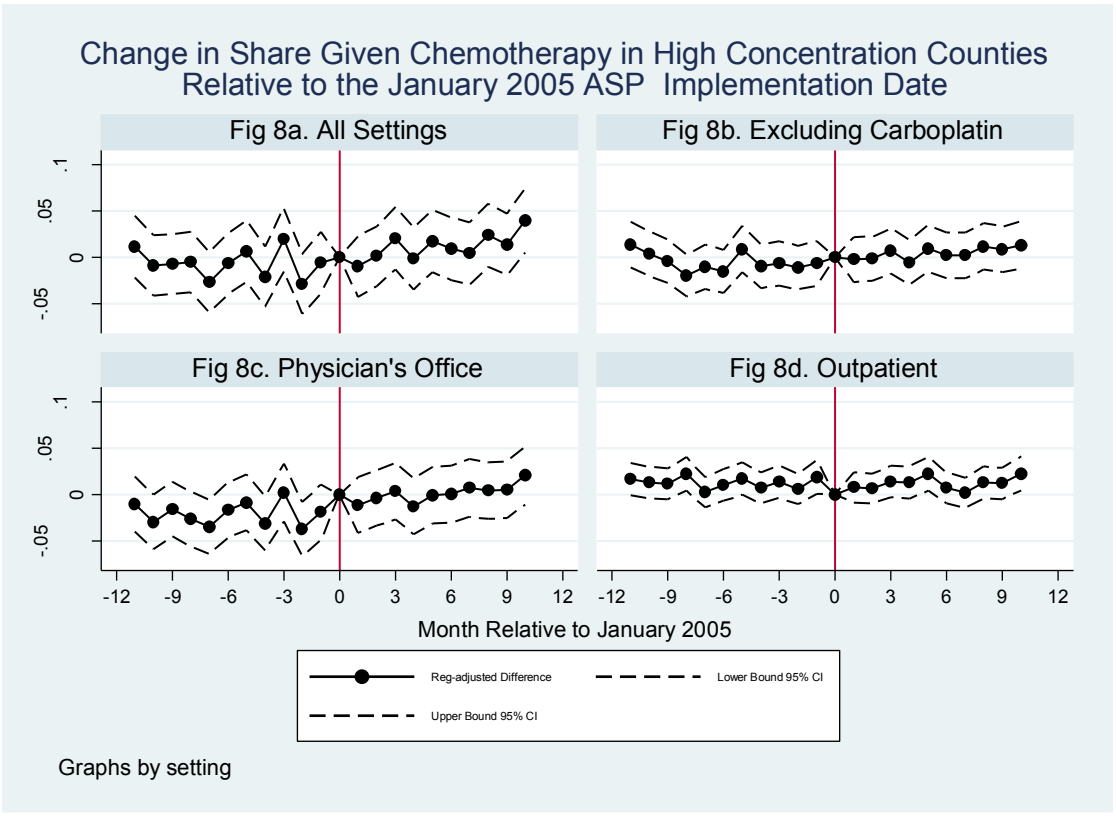

Percent Change in Chemotherapy Service Count by Month of Diagnosis
Relative to the January 2005 ASP Implementation Date
Fig 4a. All Settings
Fig 4b. Excluding Carboplatin
Fig 4c. Physician's Office
Fig 4d. Outpatient
-.2 -.1 0 .1
-12 -9 -6 -3 0 3 6 9 12
Month Relative to January 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by setting

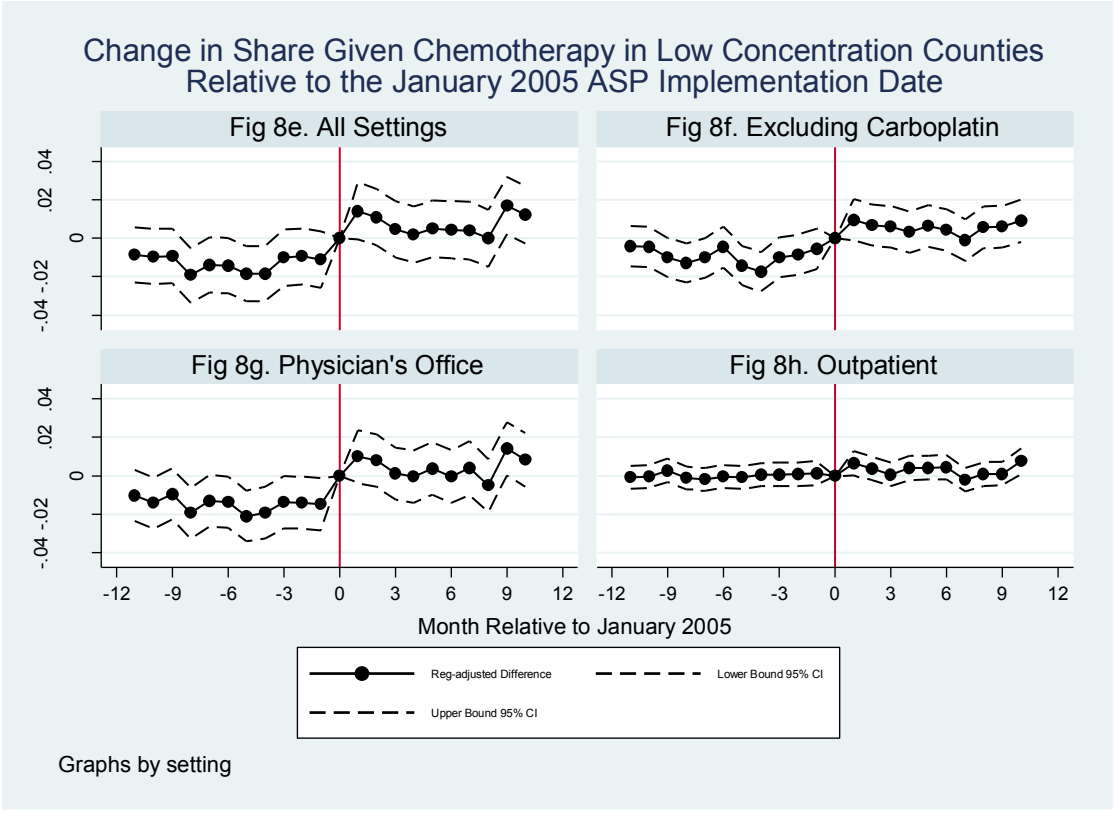

Change in the Use of Specific Agents by Month of Diagnosis
Relative to the January 2005 ASP Implementation Date
Fig 5a. Carboplatin
Fig 5b. Paclitaxel
Fig 5c. Docetaxel
Fig 5d. Etoposide
Fig 5e. Gemcitabine HCl
-.1 -.05 0 .05 .1 .15
-12 -9 -6 -3 0 3 6 9 12
Month Relative to January 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by Type

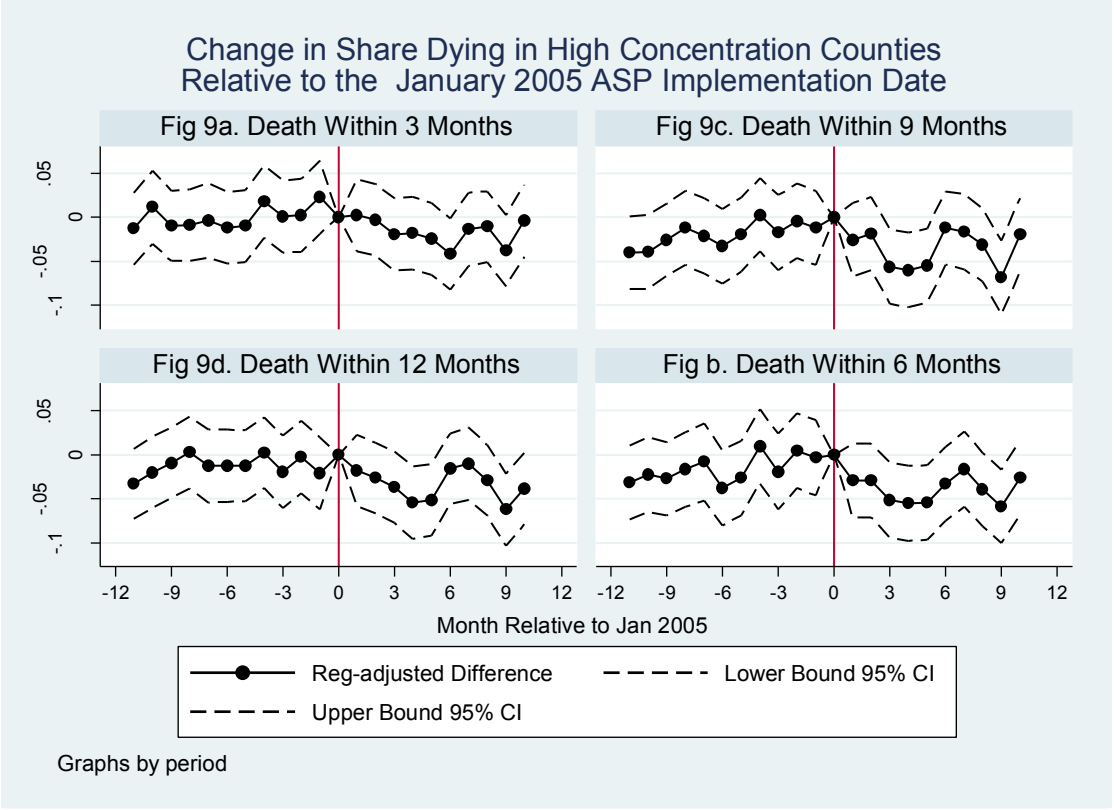

Change in the Proportion Dying by Month of Diagnosis
Relative to the January 2005 ASP Implementation Date
Fig 6a. Death Within 3 Months
Fig 6b. Death Within 6 Months
Fig 6c. Death Within 9 Months
Fig 6d. Death Within 12 Months
-.06 -.03 0 .03 .06
-12 -9 -6 -3 0 3 6 9 12
Month Relative to Jan 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by period


Change in Share Dying by Month of Diagnosis in Most Responsive States
Relative to the January 2005 ASP Implementation Date
Fig 7a. Death Within 3 Months
Fig 7b. Death Within 6 Months
Fig 7c. Death Within 9 Months
Fig 7d. Death Within 12 Months
-.06 -.03 0 .03 .06
-12 -9 -6 -3 0 3 6 9 12
Month Relative to Jan 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by period


Change in Share Dying by Month of Diagnosis in Least Responsive States
Relative to the January 2005 ASP Implementation Date
Fig 7e. Death Within 3 Months
Fig 7f. Death Within 6 Months
Fig 7g. Death Within 9 Months
Fig 7h. Death Within 12 Months
-.06 -.03 0 .03 .06
-12 -9 -6 -3 0 3 6 9 12
Month Relative to Jan 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by period


Change in Share Given Chemotherapy in High Concentration Counties
Relative to the January 2005 ASP Implementation Date
Fig 8a. All Settings
Fig 8b. Excluding Carboplatin
Fig 8c. Physician's Office
Fig 8d. Outpatient
-.05
0
.05
.1
-12
-9
-6
-3
0
3
6
9
12
Month Relative to January 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by setting


Change in Share Given Chemotherapy in Low Concentration Counties
Relative to the January 2005 ASP Implementation Date
Fig 8e. All Settings
Fig 8f. Excluding Carboplatin
Fig 8g. Physician's Office
Fig 8h. Outpatient
-.04
-.02
0
.02
.04
-12
-9
-6
-3
0
3
6
9
12
Month Relative to January 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by setting


Change in Share Dying in High Concentration Counties
Relative to the January 2005 ASP Implementation Date
Fig 9a. Death Within 3 Months
Fig 9c. Death Within 9 Months
Fig 9d. Death Within 12 Months
Fig b. Death Within 6 Months
-.1
-.05
0
.05
-12
-9
-6
-3
0
3
6
9
12
Month Relative to Jan 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by period

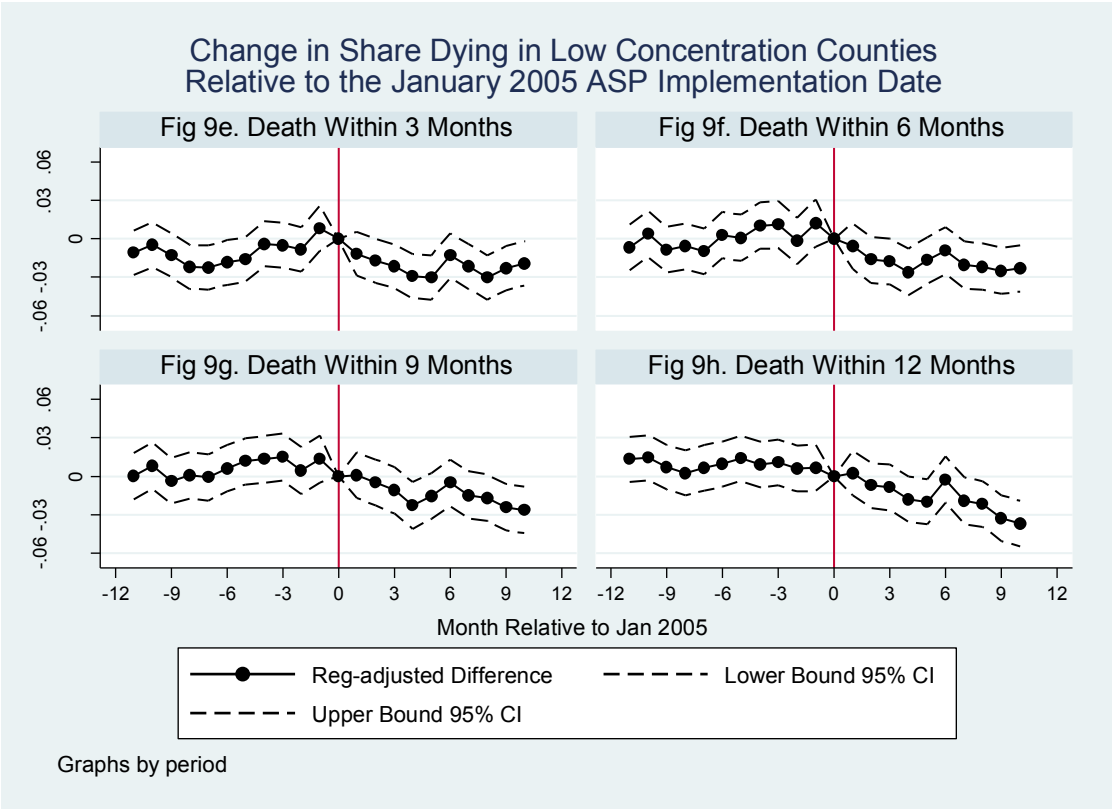

Change in Share Dying in Low Concentration Counties
Relative to the January 2005 ASP Implementation Date
Fig 9e. Death Within 3 Months
Fig 9f. Death Within 6 Months
Fig 9g. Death Within 9 Months
Fig 9h. Death Within 12 Months
-.06
-.03
0
.03
.06
-12
-9
-6
-3
0
3
6
9
12
Month Relative to Jan 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by period

Table 1. Sample Characteristics overall and in the11 Months Pre versus Post Jan 2005 Reform*

	Overall N=132,768	Pre N=66,501	Post N=66,267
Age at Diagnosis – mean (sd)	74.0 (8.35)	74.0 (8.32)	74.1 (8.38)
Share Male	0.513	0.516	0.510
Distribution by Race/ethnicity			
White	0.877	0.876	0.877
African-American	0.089	0.090	0.088
Asian	0.0108	0.0105	0.011
Other/Unknown	0.0091	0.0087	0.0095
Hispanic	0.0102	0.0104	0.0099
Native American	0.0039	0.0038	0.0039
Deyo-Charlson Score – mean (sd)	1.07 (1.30)	1.06 (1.30)	1.08 (1.31)
Metastasis within 1 month	0.289	0.289	0.289
Quarterly Reimbursement Rates Per Standard Monthly Dose ($)			
carboplatin	1540 (153)	1845 (29)	930 (261)
paclitaxel	1590 (294)	2272 (70.9)	225 (9.28)
docetaxel	2657 (58.4)	2732 (74.8)	2506 (6.87)
etoposide	77.6 (27.8)	111(36.8)	11.4(0.369)
gemcitabine	1311 (13.5)	1313 (20.6)	1305 (1.71)

* The pre-period is Feb-Dec 2004 and the post-period is Jan-Nov 2005. Standard deviations are in parentheses. Percentages may not sum to one due to rounding. Reimbursement rates are from the CMS website. The sample includes Medicare beneficiaries diagnosed with lung cancer from Feb 2004 to Nov 2005. Chemotherapy "excluding carboplatin" means the share receiving chemotherapy but not carboplatin. The Deyo-Charlson score is a comorbidity index measured in the year prior to a cancer diagnosis.

Table 2. Summary Outcomes Overall and in the 11 Months Pre versus Post Reform*

	Overall N=132,768	Pre N=66,501	Post N=66,267
Share of Patients Receiving Chemotherapy Within …			
1 month	0.175	0.166	0.185
1 month, excluding carboplatin	0.080	0.073	0.087
1 month, physician's office	0.140	0.130	0.149
1 month, outpatient hospital	0.0285	0.0295	0.0275
3 months	0.279	0.263	0.295
3 months, excluding carboplatin	0.103	0.093	0.113
3 months, physician's office	0.228	0.213	0.244
3 months, outpatient hospital	0.067	0.065	0.069
Chemotherapy Drugs Conditional on Treatment within 1 month			
Share carboplatin	0.543	0.557	0.531
Share paclitaxel	0.285	0.315	0.257
Share docetaxel	0.090	0.083	0.096
Share etoposide	0.208	0.211	0.206
Share gemcitabine	0.092	0.100	0.085
Patients Dying Within…			
3 months	0.333	0.337	0.328
6 months	0.458	0.467	0.449
9 months	0.516	0.525	0.507
12 months	0.605	0.616	0.594
No date of death	0.110	0.100	0.120

* The pre-period is Feb-Dec 2004 and the post-period is Jan-Nov 2005. Standard deviations are in parentheses. Percentages may not sum to one due to rounding. Reimbursement rates are from the CMS website. The sample includes Medicare beneficiaries diagnosed with lung cancer from Feb 2004 to Nov 2005. Chemotherapy "excluding carboplatin" means the share receiving chemotherapy but not carboplatin. The Deyo-Charlson score is a comorbidity index measured in the year prior to a cancer diagnosis.

Table 3. Changes in Chemotherapy Treatment After the January 2005 Payment Change

	Before Jan/05	Feb/04 –Nov/05	Apr/04 –Sept/05
Share of Patients Receiving Chemotherapy Treatment…			
within 1 month of Diagnosis	0.166	0.015	0.014
		(0.003)	(0.002)
within 1 month, excluding carboplatin	0.073	0.014	0.012
		(0.002)	(0.002)
within 1 month, in a Physician's Office	0.130	0.016	0.016
		(0.003)	(0.003)
within 1 month, in an outpatient hospital clinic	0.028	-0.001	-0.001
		(0.001)	(0.001)
Number of Observations (weeks)	48	96	72

Notes: Means in column 1 are for the dependent variable prior to ASP implementation, Feb-Dec 2004. Columns 2 - 3 present estimates from separate time-series regressions. Estimates are the coefficients on an indicator for the pot-reform/ASP payment period. Regressions control for mean patient characteristics (see Table 1) and relative-week trends that are allowed to differ on either side of the payment change. Excluding carboplatin" means the share of patients treated with chemotherapy and without carboplatin. Newey-West standard errors allowing for autocorrelation up to 52-week lags are given in parentheses.

Table 4. Change in Chemotherapy Service Counts after the January 2005 Payment Change

	Before Jan/05	Feb/04 –Nov/05	Apr/04 –Sept/05
Percent Change in Monthly Chemotherapy Service…			
within 1 month of Diagnosis	1.00	0.339	0331
		(0.046)	(0.044)
within 1 month, excluding carboplatin	0.304	0.203	0.209
		(0.025)	(0.022)
within 1 month, in a Physician's Office	0.916	0.347	0.340
		(0.047)	(0.046)
within 1 month, in an outpatient hospital clinic	0.087	0.013	0.013
		(0.003)	(0.004)
Number of Observations (weeks)	48	96	72

Notes: See notes to Table 3.

Table 5. Changes in Types of Chemotherapy Drugs

	Before Jan/05	Feb/04 –Nov/05	Apr/04 –Sept/05
Share of Patients Receiving…			
carboplatin	0.092	0.001	0.0017
		(0.001)	(0.0017)
paclitaxel	0.053	-0.0053	-0.0051
		(0.0014)	(0.002)
docetaxel	0.014	0.0043	0.0037
		(0.002)	(0.002)
etoposide	0.035	0.0032	0.0044
		(0.001)	(0.001)
gemcitabine	0.017	0.0002	0.001
		(0.0008)	(0.0007)
Share of Chemotherapy-treated Patients Receiving…			
carboplatin	0.557	-0.039	-0.026
		(0.008)	(0.008)
paclitaxel	0.300	-0.074	-0.061
		(0.007)	(0.006)
docetaxel	0.083	0.026	0.026
		(0.004)	(0.006)
etoposide	0.210	-0.005	0.004
		(0.010)	(0.007)
gemcitabine	0.100	0.007	0.008
		(0.003)	(0.004)
Number of Observations (weeks)	48	96	72

Notes: See notes to Table 3

Table 6. Change in the Share of Patients Dying within 3, 6, 9 or 12 Months of Diagnosis

	Mean Before Jan/05	Feb/04 –Nov/05	Apr/04 –Sept/05
Share of Patients Dying within…			
3 months of diagnosis	0.339	-0.009 (0.004)	-0.010 (0.005)
6 months of diagnosis	0.469	-0.016 (0.002)	-0.016 (0.004)
9 months of diagnosis	0.527	-0.011 (0.002)	-0.011 (0.004)
1 year of diagnosis	0. 619	-0.001 (0.003)	-0.001 (0.003)
Number of Observations (weeks)	48	96	72

Notes: see notes to Tables 3.

Table 7. Change in the Share of Patients Dying: Above vs. Below Median Responsive States

	Above Median Response			Below Median Response		
	Mean	Feb/04 –Nov/05	Apr/04 –Sept/05	Mean	Feb/04 –Nov/05	Apr/04 –Sept/05
Change in Share Dying within…						
3 months	0.338	-0.011 (0.006)	-0.023 (0.003)	0.340	0.014 (0.002)	0.014 (0.003)
6 months	0.466	-0.015 (0.004)	-0.023 (0.004)	0.473	0.003 (0.005)	0.006 (0.007)
9 months	0.524	-0.008 (0.004)	-0.017 (0.004)	0.531	-0.0001 (0.006)	0.003 (0.007)
1 year	0. 615	-0.0001 (0.003)	-0.007 (0.004)	0.623	0.009 (0.007)	0.007 (0.008)
Obs (weeks)	48	96	72	48	96	72

Notes: Means in column 1 and 4 are for the dependent variable for states with above or below median changes in chemotherapy treatment in the period prior to the reform. See notes to Tables 3 for other details.

Table 8: Change in Treatment and Survival: Above vs. Below Median Age Patients

	Above Median Age Patients			Below Median Age Patients		
	Mean	Feb/04 –Nov/05	Apr/04 –Sept/05	Mean	Feb/04 –Nov/05	Apr/04 –Sept/05
Share of Patients Receiving Chemotherapy Treatment within…						
1 month of diagnosis	0.119	0.014 (0.002)	0.014 (0.003)	0.210	0.024 (0.003)	0.020 (.003)
Share of Patients Dying Within…						
3 months of diagnosis	0.397	-0.023 (0.006)	-0.024 (0.005)	0.285	-0.0059 (0.0035)	-0.0063 (0.0036)
6 months of diagnosis	0.530	-0.029 (0.004)	-0.032 (0.005)	0.411	-0.015 (0.002)	-0.020 (0.0023)
9 months of diagnosis	0.588	-0.026 (0.003)	-0.031 (0.004)	0.470	-0.012 (0.003)	-0.017 (0.003)
1 year of diagnosis	0. 675	-0.015 (0.002)	-0.018 (0.003)	0.567	0.002 (0.003)	-0.006 (0.0038)
Observations (weeks)	48	96	72	48	96	72

Notes: Means in column 1 and 4 are for the dependent variable for patients above versus below the median age patient at diagnosis in the period prior to the reform, Feb-Nov 2004. See notes to Table 3 for other details.

Table 9. Change in Treatment and Survival: Counties with a High vs. Low Concentration of Providers Administering Chemotherapy Treatment

	High Concentration Counties			Low Concentration Counties		
	Mean	Feb/04 –Nov/05	Apr/04 –Sept/05	Mean Pre-Jan	Feb/04 –Nov/05	Apr/04 –Sept/05
Chemotherapy Treatment within 1 month…						
Share Receiving Any	0.169	-0.0004 (0.005)	0.001 (0.007)	0.165	0.020 (0.002)	0.019 (0.003)
Percent Change in Services	0.973	0.203 (0.045)	0.185 (0.045)	1.01	0.389 (0.031)	0390 (0.030)
Share Dying Within…						
3 months of diagnosis	0.351	-0.017 (0.006)	-0.010 (0.004)	0.337	-0.006 (0.0036)	-0.008 (0.004)
6 months of diagnosis	0.482	-0.028 (0.008)	-0.022 (0.006)	0.466	-0.012 (0.002)	-0.014 (0.002)
9 months of diagnosis	0.541	-0.027 (0.008)	-0.019 (0.005)	0.524	-0.004 (0.003)	-0.007 (0.002)
1 year of diagnosis	0. 635	-0.008 (0.008)	0.0003 (0.006)	0.616	0.006 (0.005)	0.0001 (0.002)
Observations (weeks)	48	96	72	48	96	72

Notes: Means in column 1 and 3 are for the dependent variable for patients in counties with low vs. high concentration in the period prior to the reform, Feb-Nov 2004. The mean for service counts is without logs. High concentration is defined as an HHI greater than 0.25. See notes to Table 3 for other details.

Appendix

Proof of Proposition 1

Proof. Let $s_m^j = q_m^j/\eta_j$ the change in the share of patients treated by physician j in response to a change in m. Then for $\alpha \approx 0$, equation 3 can be rewritten as

$$s_m^j \cong \frac{V_\pi + V_{\pi\pi}\pi}{\eta_j e_{qq}}. \tag{6}$$

Denote the inverse of the change in the share of patients treated as $h(\eta_j) = ||\frac{1}{s_m^j}|| = \eta_j\kappa$ where $\kappa \equiv ||\frac{e_{qq}}{V_\pi + V_{\pi\pi}\pi}||$. The market level measure of h is then given by the weighed averages of the individual $h(\eta_j)$:

$$H_m = \Sigma_{j=1}^J \eta_j h(\eta_j). \tag{7}$$

An internal solution for equation 2 implies that κ is invariant with respect to η, so H_m can be rewritten as

$$H_m = \Sigma_{j=1}^J \eta_j \eta_j \kappa = \kappa \Sigma_{j=1}^J \eta_j \eta_j. \tag{8}$$

So the change in the share of patients treated in a market m is equivalent to a scaled version of the Herfindahl-Hirschman Index (HHI). And since HHI belongs to the family of "allowable" concentration indices as described in Encaoua and Jacquemin (1980), H_m will be positively correlated to any other such allowable concentration index (e.g. the entropy index).

Then since H_m is positively correlated to market concentration, its inverse (i.e. the magnitude of the change in the share of patients treated in a market) will be negatively correlated with measures of market concentration. ■

Proof of Proposition 2

Proof. The change in patient welfare generated by a physician j's response to a change in profit margin m is given by $b(q^j/\eta_j)q_m^j$. The magnitude of the market level change in patient welfare can thus be written as $W_m = \Sigma_{j=1}^J \eta_j ||b(q^j/\eta_j)q_m^j||$, which

for small $\alpha \approx 0$ can be rewritten as

$$W_m \cong \Sigma_{j=1}^{J} \eta_j h(\eta_j) \text{ where } h(\eta_j) = b(q^j/\eta_j)||\tfrac{V_\pi + V_{\pi\pi}\pi}{e_{qq}}||. \tag{9}$$

Encaoua and Jacquemin (1980) show that for a concentration index defined as

$$\zeta(\alpha_1,', \alpha_N) = \Sigma_{i=1}^{N} \alpha_i g(\alpha_i), \tag{10}$$

where the α_is represents the market share of firm i, if $g(\alpha)$ is any arbitrary function that is weakly increasing in α, if $\alpha g(\alpha)$ is convex in α, it belongs to a family of "allowable" concentration indices. Such allowable indices (which include HHI and the entropy index) satisfies the following properties: it is invariant to permutations of market shares between firms i, it satisfies the Lorentz condition that a mean pressuring spread increases ε, and the concentration of symmetric firms decreases in the number of firms. Thus it is sufficient to show that $h(\eta)$ is increasing in η and $\eta h(\eta)$ is convex in η to prove the proposition.

To show $h(\eta)$ is weakly increasing in η, we take the derivative of h with respect to η and get

$$\frac{dh}{d\eta} = -\frac{q^j}{\eta^2} b' || \frac{V_\pi + V_{\pi\pi}\pi}{e_{qq}} ||. \tag{11}$$

Then since $b' \leq 0$, equation 11 is positive and $h(\eta)$ is weakly increasing in η.

To show convexity of $\eta h(\eta)$ we take the second derivative and get (after some simplification) the expression

$$\frac{d^2(\eta h)}{d\eta^2} = \frac{(q^j)^2}{\eta^3} b'' || \frac{V_\pi + V_{\pi\pi}\pi}{e_{qq}} ||. \tag{12}$$

Since by assumption $b'' > 0$, $\frac{d^2(\eta h)}{d\eta^2} > 0$. ■

Two-Payer Model

In this section, we analyze the model presented in the paper, with the addition of a second, private payer. Here we show that the two propositions in the paper hold in a

two payer model, and generate two addition predictions regarding how the share for public vs. private patients affects physician response to a fee cut.

In addition the assumptions of the main model, we incorporate a second payer for medical care. In each market k, for patients of type b, a fraction γ are public (i.e. Medicare) patients while the remaining $(1-\gamma)$ are private payers. Notationally, we denote the number of public patients treated by q and the number of private payer patients treated by q'. Similarly the profit margin for treating each type of patient is denoted by m and m' respectively.

Physician utility function can then be written as

$$U(q) = V(\pi) - e(q+q') + \alpha\eta[\gamma \int_0^{q/\eta\gamma} b(x)dx + (1-\gamma)\int_0^{q'/\eta(1-\gamma)} b(x)dx], \quad (13)$$

where $\pi = mq + m'q'$ and α represents the weight physicians place on patient benefit. The FOCs for utility maximization are then given by

$$\begin{aligned} mV_\pi - e_q + \alpha b(q/\eta\gamma) &= 0 \quad (14) \\ m'V_\pi - e_q + \alpha b(q'/\eta(1-\gamma)) &= 0. \quad (15) \end{aligned}$$

As before concern for patient welfare (αb) pushes physicians towards providing what, from the patient's perspective, is the optimal level of care q^P, but in general physicians may either under or over provide care relative to this optimum: in the limit $\alpha \to \infty$, $q^* \to q^P$.

To determine physician response in q to a change in profit margin m, we take the derivative of equation 14 with respect to m. Rearranging, we get the following relationship:

$$q_m = \frac{V_\pi + V_{\pi\pi}\pi + q'_m(V_{\pi\pi}mm' - e_{qq})}{e_{qq} + \frac{\alpha}{\eta\gamma}b_q} \quad (16)$$

As before q_m can be positive or negative. That is, it cannot be determined ex-ante whether an increase in profitability leads to an increase or decrease in provision of the service experiencing a price cut. But now q_m is driven by the tradeoff between decreasing marginal returns (the substitution effect) and the increasing marginal returns to income (the income effect), as well as "cross-service effects" (i.e. changes in

the marginal cost of effort of q due to changes in q').

Note that in the case where $q'_m \approx 0$, equation 16 is identical to equation 3, with $\eta\gamma$, the Medicare market share, taking taking the place of overall market share η. This is the case described in McGuire and Pauly (1991): "When income effects are absent and either the marginal utility of leisure is unchanged... we get a surprising conclusing: there will be no effects" on the quantity of service provided to the private payer in response to a change in public payer margins. In such a case the "bite of a price cut will decline with market share". That is doctors in markets with higher share of Medicare patients will be more affected by the price cut than those in low Medicare patient share markets.

In this case, in addition to proposition 1 and 2, and we get the following corollaries:

Corollary 3 *If Proposition 3 holds, H_m^k is increasing in γ.*

Proof. Since $\frac{d\eta}{d\gamma} = \frac{d\kappa}{d\gamma} = 0$, $\frac{dH_m^k}{d\gamma} = \kappa\Sigma_{j=1}^J \eta_j\eta_j > 0$ ■

Corollary 4 *If Proposition 4 holds, W_m^k is increasing in γ.*

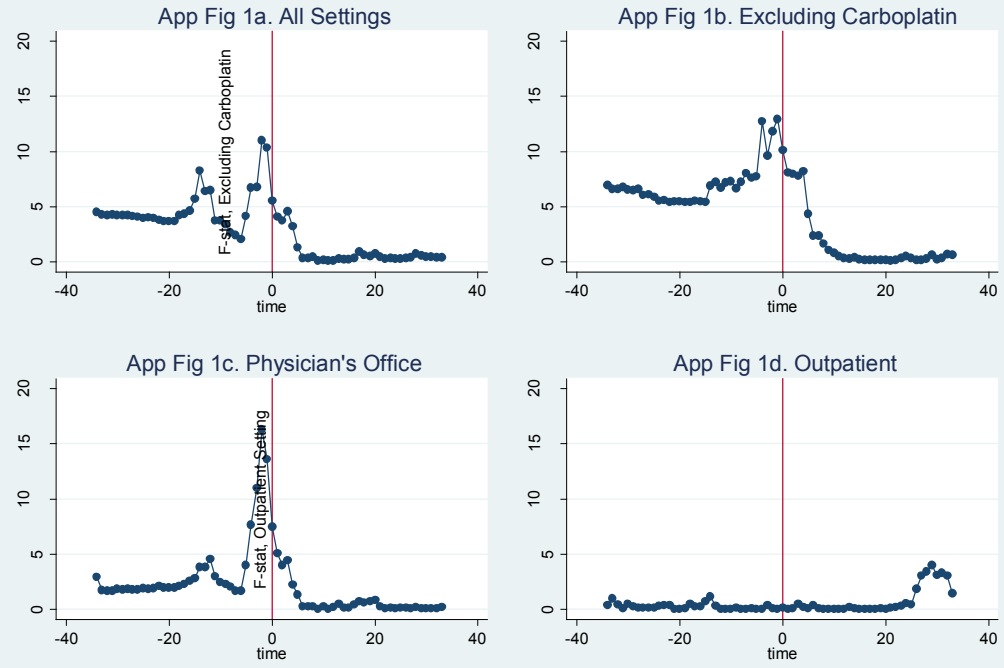
Proof. $\frac{dW_m^k}{d\gamma_k} = \Sigma_{j=1}^J(\frac{-q^j}{\gamma^2})b'||\frac{V_\pi + V_{\pi\pi}\pi + q'_m(V_{\pi\pi}mm' - e_{qq})}{e_{qq}}||$. Since $b' < 0$, $\frac{dW_m^k}{d\gamma_k} < 0$ ■

These corollaries simply state that when the the share of public patients is larger, a cut in public patient margins will have a larger impact on physician behavior.

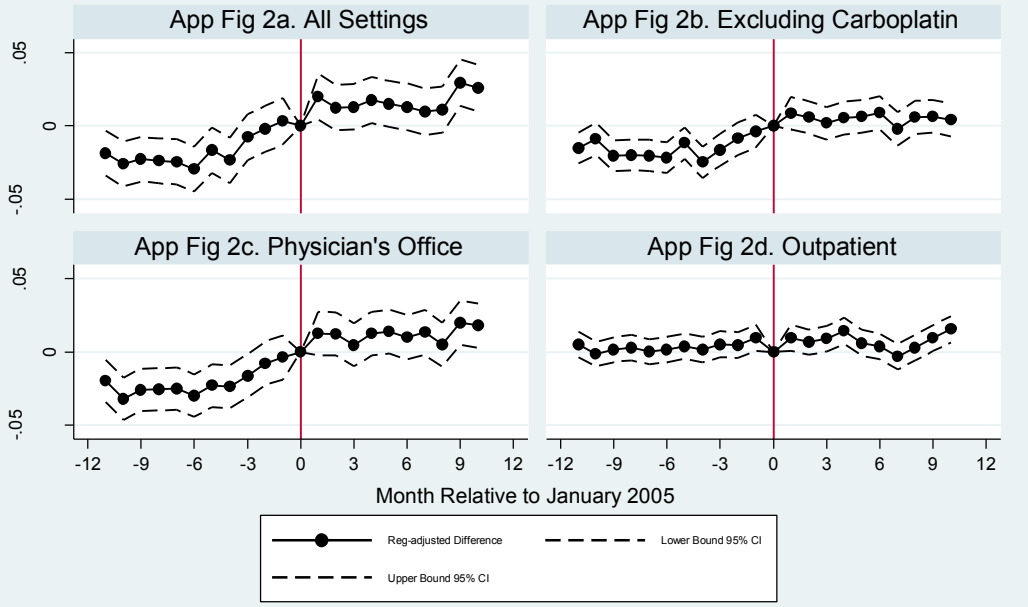
In the case where there are significant "cross service effects" (i.e. $||q'_m|| >> 0$), the expressions that describe physician response are quite complex, and will depend on the specific values of the various parameters. In such a case, while proposition 1 and 2 hold, the two corollaries need not, and will depend on the specific form of the utility function, as well as the parameter values.

While the predictions regarding share Medicare patient share and physician response are straightforward, data on the share of Medicare patients vs. private payer patients is largely unavailable, and so we were unable to empirically test these additional predictions in this paper.

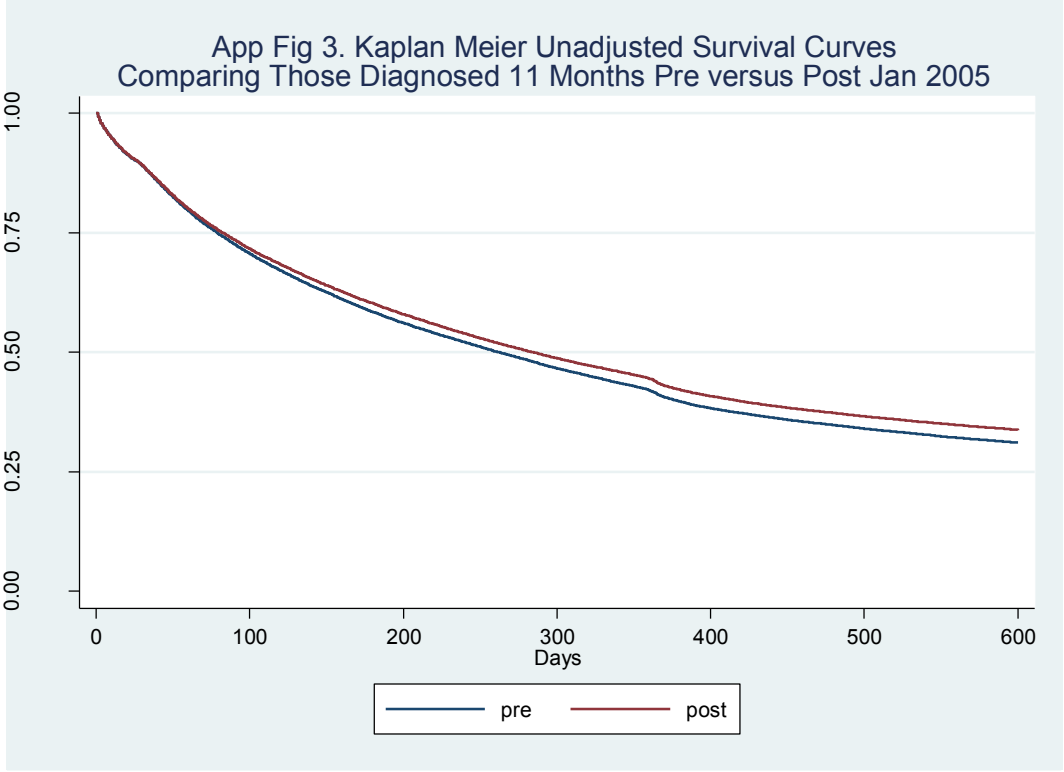
Appendix Figure 1: F-statistics Testing for a Break in Treatment



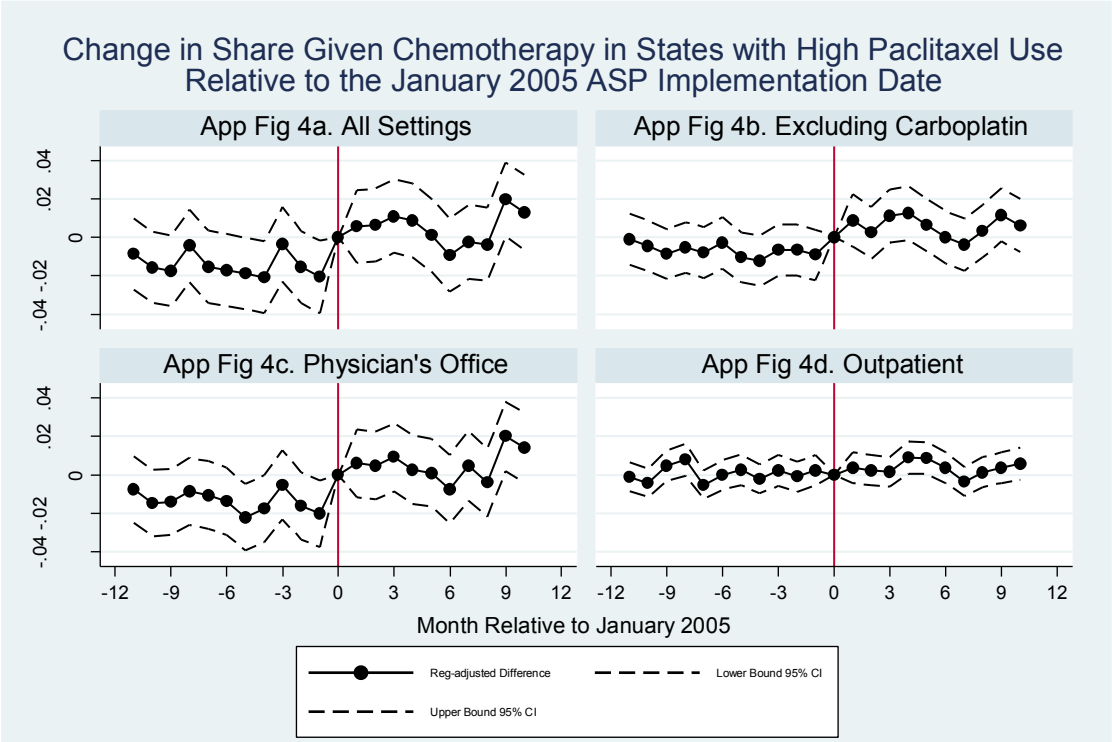
Change in the Share Given Chemotherapy Within 90 Days By Month of Diagnosis Relative to the January 2005 ASP Implementation Date

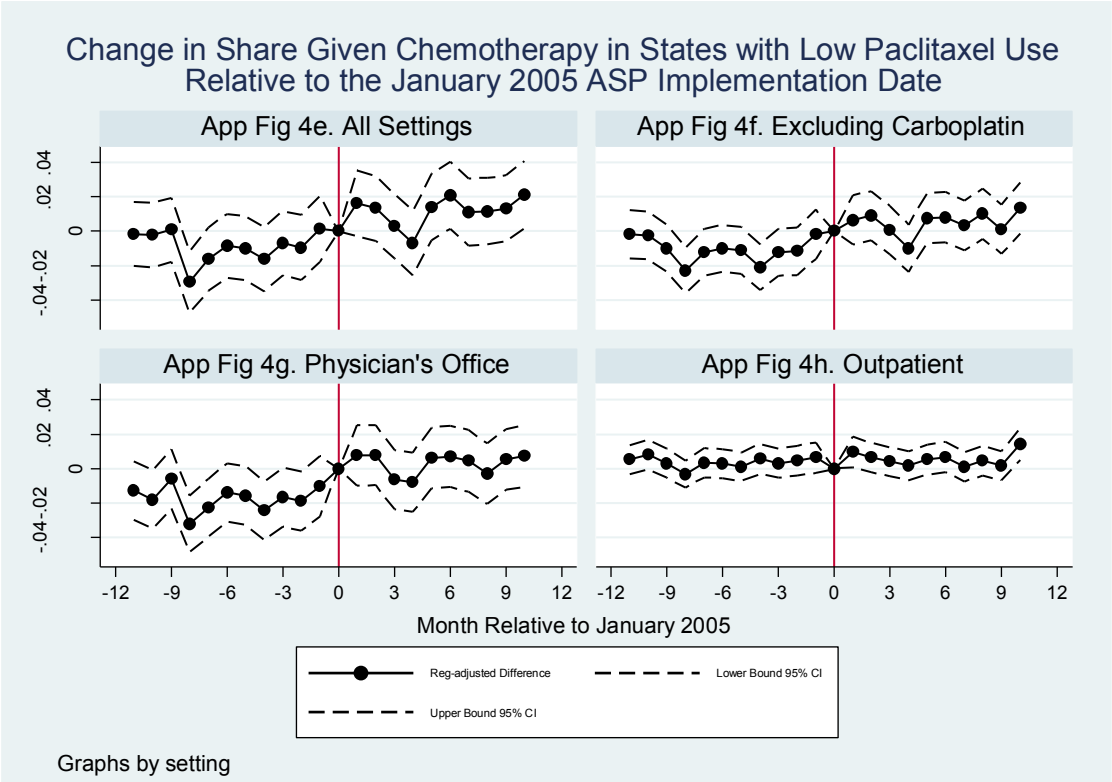


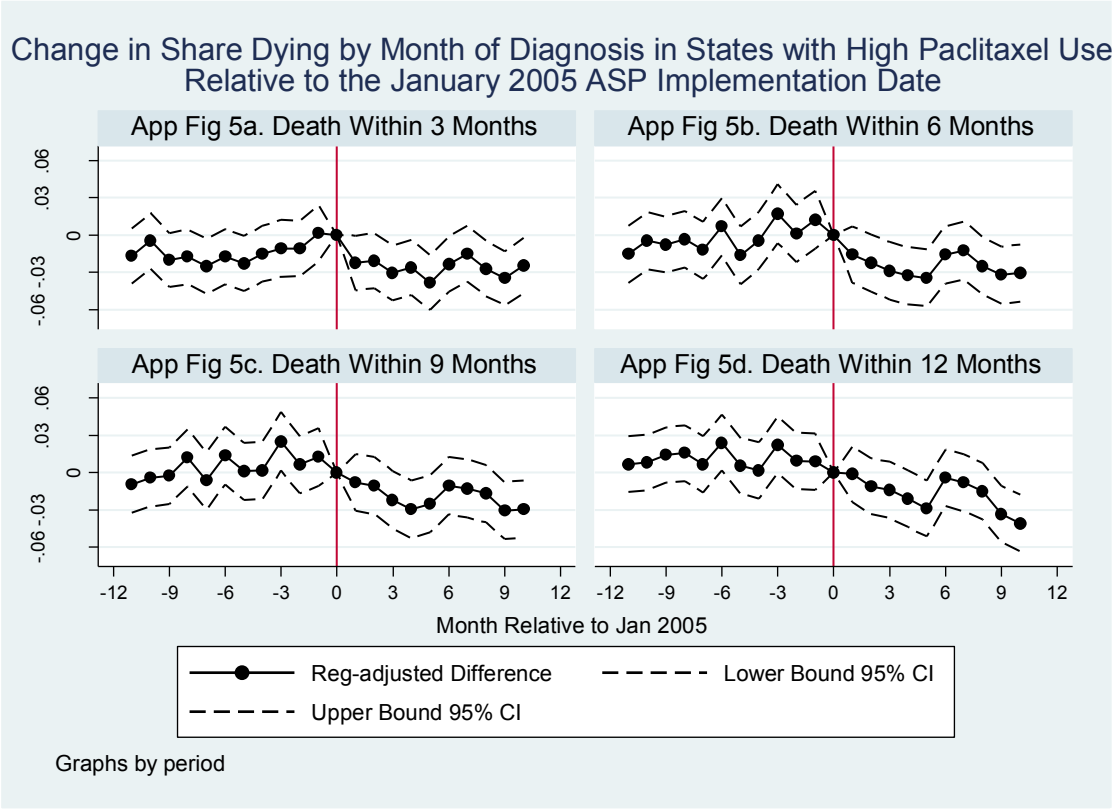
Graphs by setting

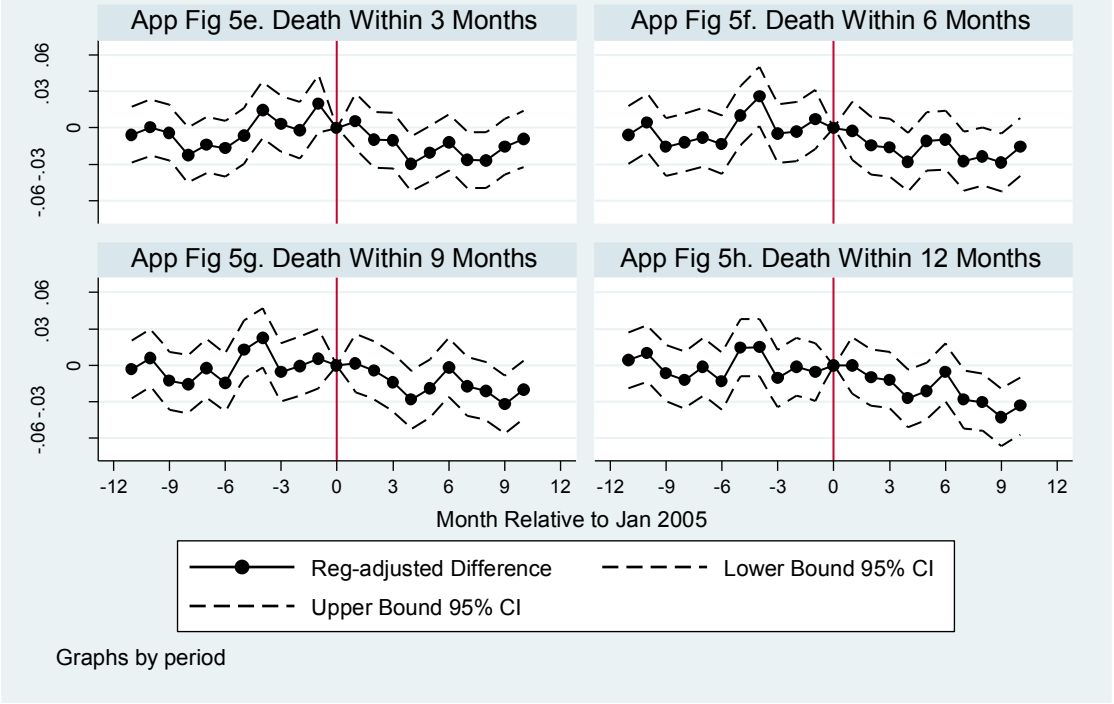


Note: There is a small discontinuity we cannot explain at about 365 days in both the cohorts diagnosed in the pre and post periods. Since it is of similar magnitude across cohorts, it should not affect our results.

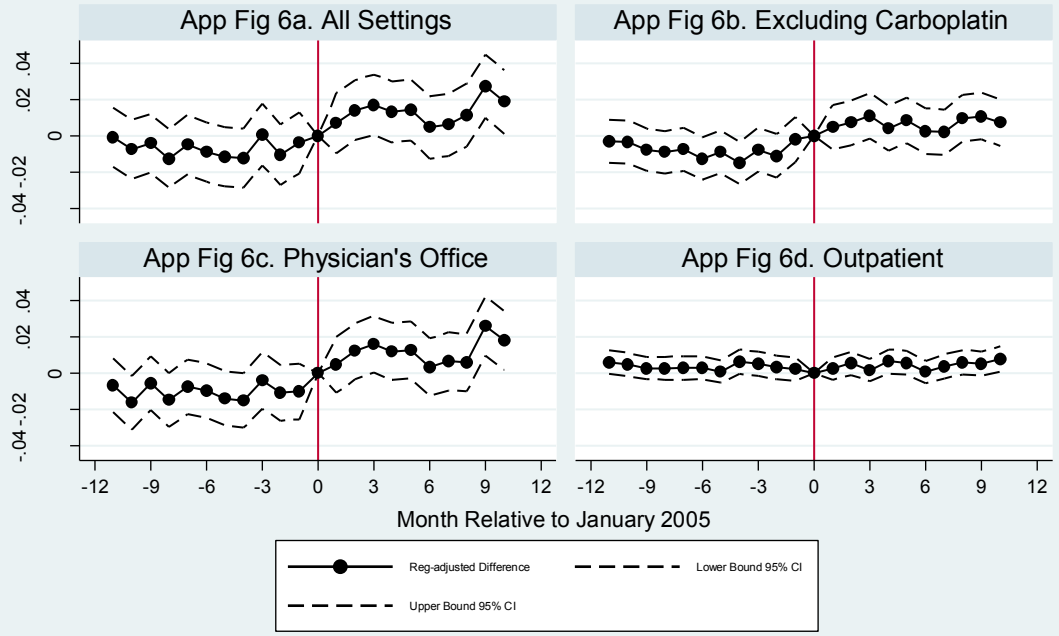

Change in Share Given Chemotherapy in States with High Paclitaxel Use Relative to the January 2005 ASP Implementation Date
App Fig 4a. All Settings
App Fig 4b. Excluding Carboplatin
App Fig 4c. Physician's Office
App Fig 4d. Outpatient
-.04 -.02 0 .02 .04
-12 -9 -6 -3 0 3 6 9 12
Month Relative to January 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by setting


Change in Share Given Chemotherapy in States with Low Paclitaxel Use Relative to the January 2005 ASP Implementation Date
App Fig 4e. All Settings
App Fig 4f. Excluding Carboplatin
App Fig 4g. Physician's Office
App Fig 4h. Outpatient
-.04 -.02 0 .02 .04
-12 -9 -6 -3 0 3 6 9 12
Month Relative to January 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by setting


Change in Share Dying by Month of Diagnosis in States with High Paclitaxel Use Relative to the January 2005 ASP Implementation Date
App Fig 5a. Death Within 3 Months
App Fig 5b. Death Within 6 Months
App Fig 5c. Death Within 9 Months
App Fig 5d. Death Within 12 Months
-.06 -.03 0 .03 .06
-12 -9 -6 -3 0 3 6 9 12
Month Relative to Jan 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by period

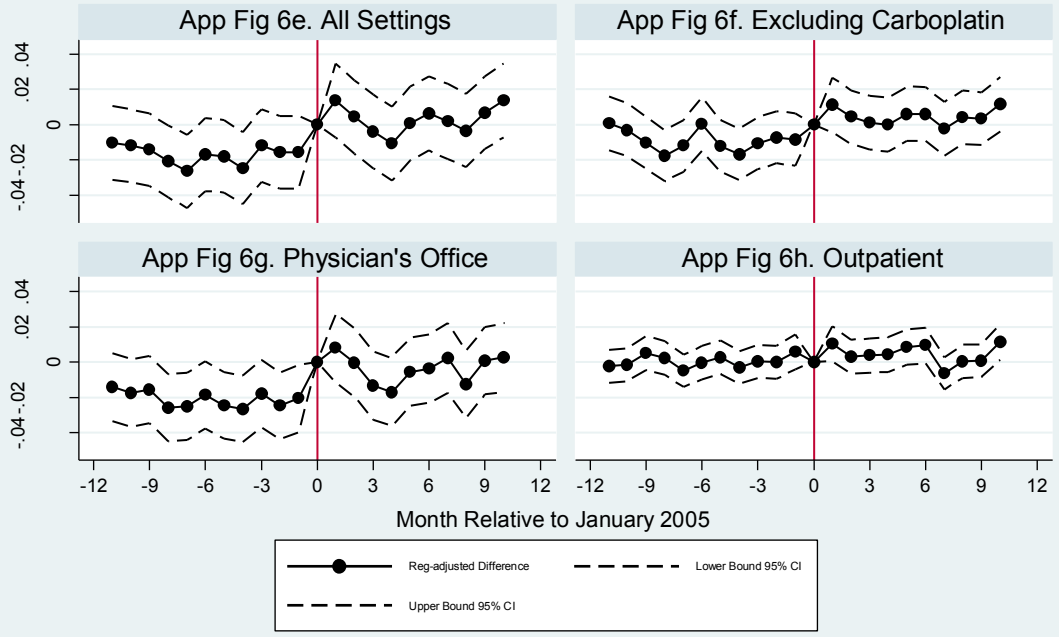

Change in Share Dying by Month of Diagnosis in States with Low Paclitaxel Use Relative to the January 2005 ASP Implementation Date
App Fig 5e. Death Within 3 Months
App Fig 5f. Death Within 6 Months
App Fig 5g. Death Within 9 Months
App Fig 5h. Death Within 12 Months
-.06 -.03 0 .03 .06
-12 -9 -6 -3 0 3 6 9 12
Month Relative to Jan 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by period

Change in Share Given Chemotherapy Among Above Median Age Patients Relative to the January 2005 ASP Implementation Date

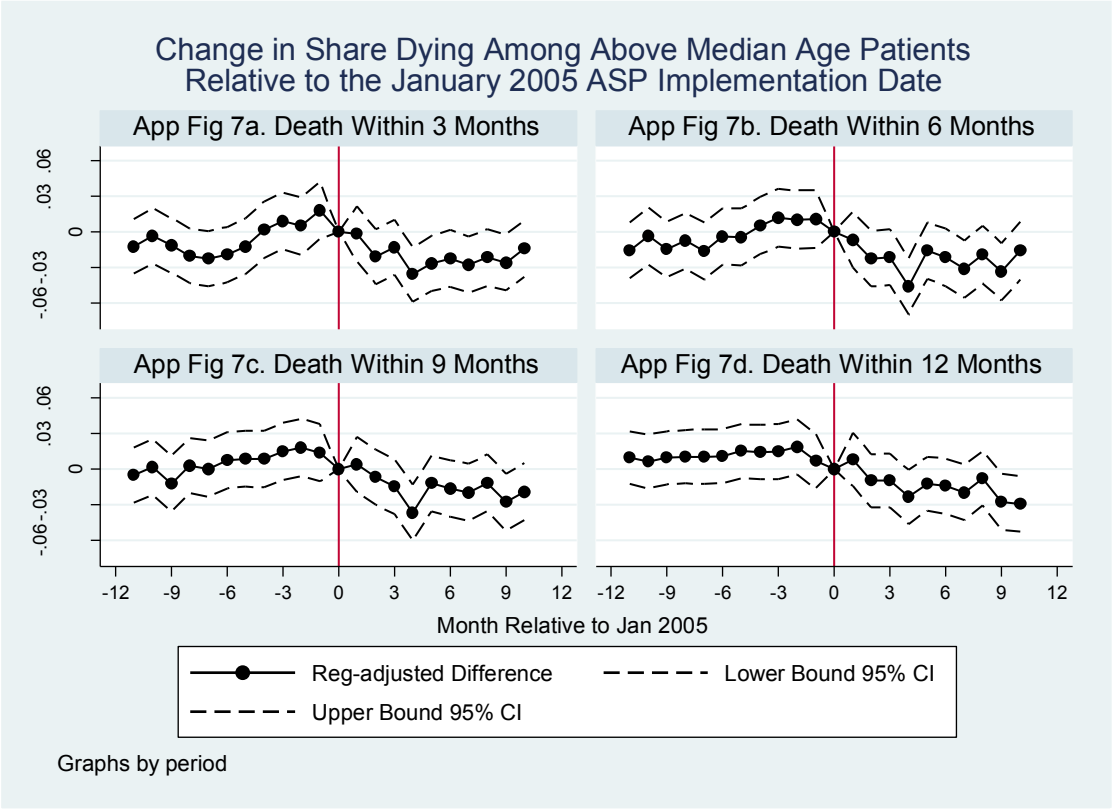


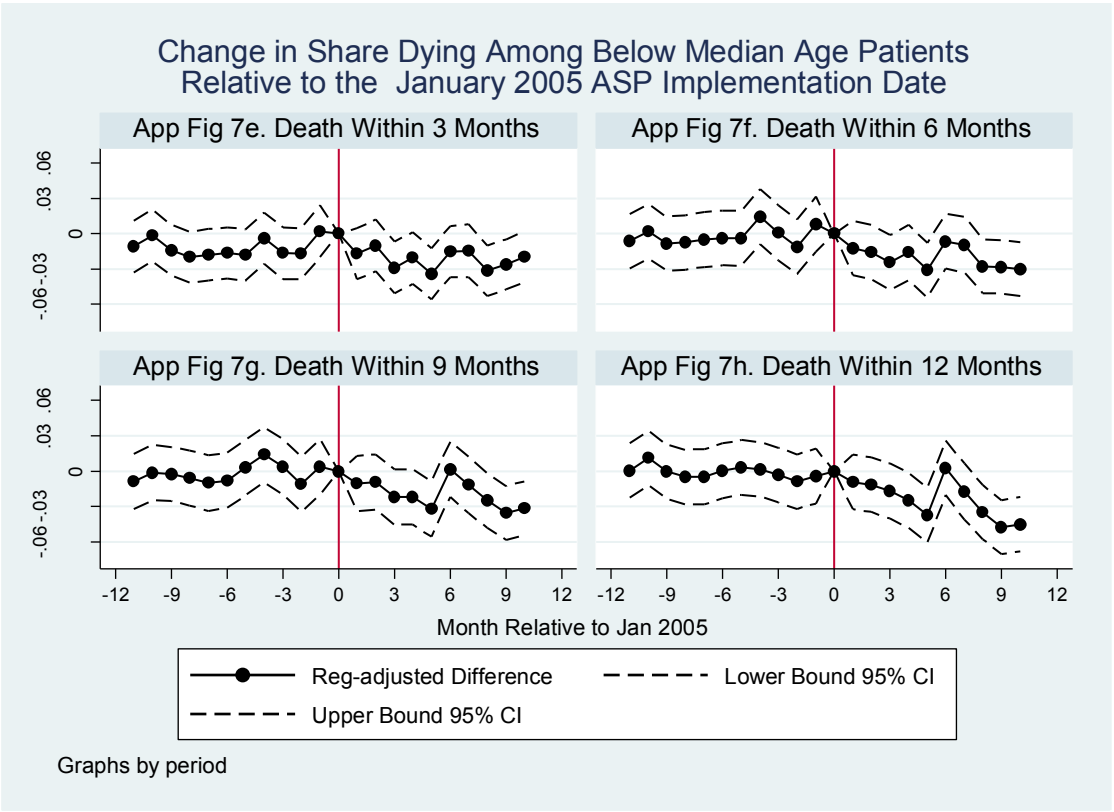
Graphs by setting

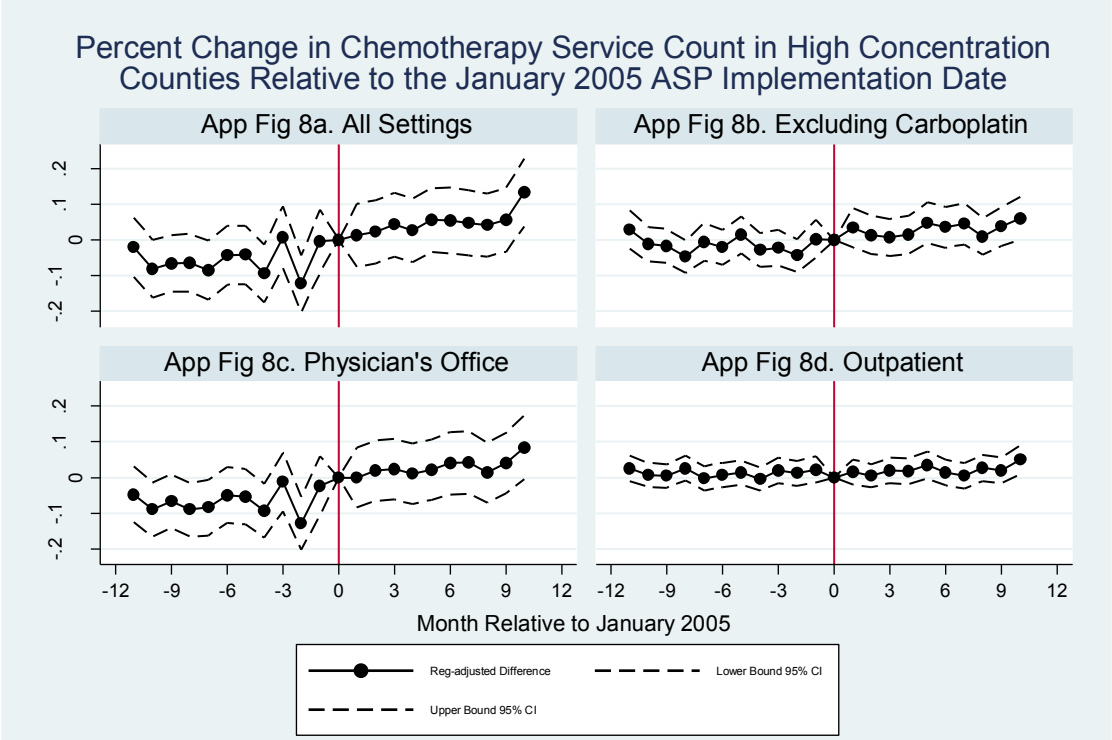
Change in the Share Given Chemotherapy Among Below Median Age Patients Relative to the January 2005 ASP Implementation Date

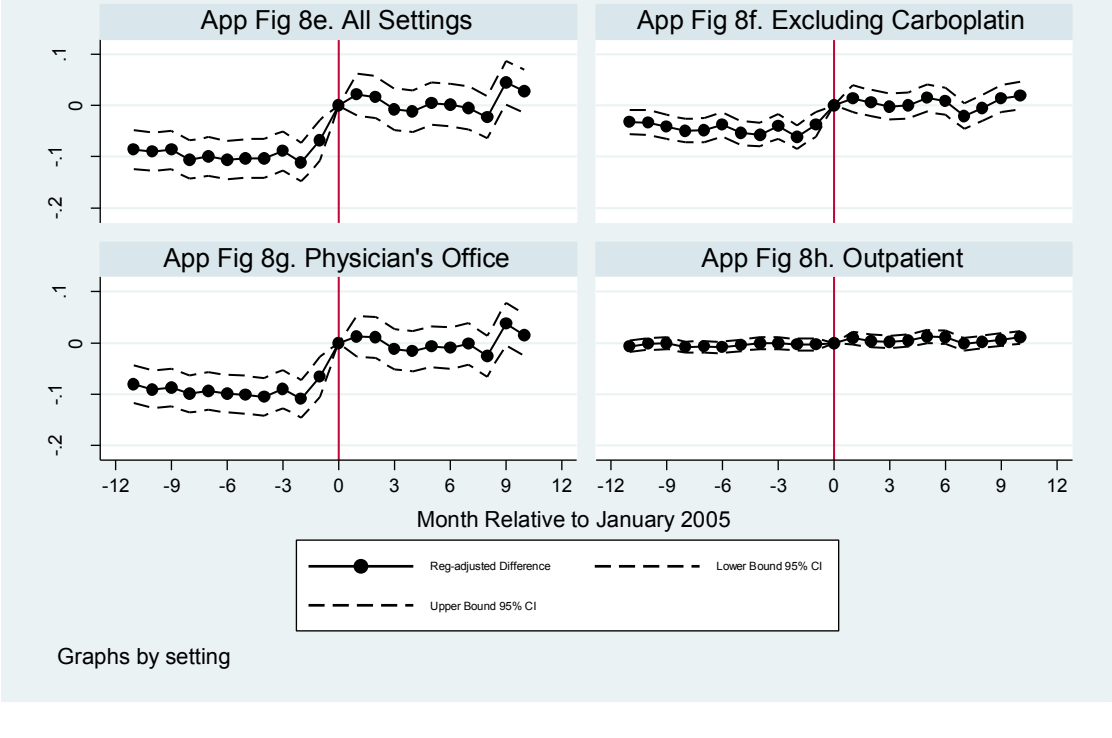


Graphs by setting


Change in Share Dying Among Above Median Age Patients
Relative to the January 2005 ASP Implementation Date
App Fig 7a. Death Within 3 Months
App Fig 7b. Death Within 6 Months
App Fig 7c. Death Within 9 Months
App Fig 7d. Death Within 12 Months
-.06 -.03 0 .03 .06
-12 -9 -6 -3 0 3 6 9 12
Month Relative to Jan 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by period


Change in Share Dying Among Below Median Age Patients
Relative to the January 2005 ASP Implementation Date
App Fig 7e. Death Within 3 Months
App Fig 7f. Death Within 6 Months
App Fig 7g. Death Within 9 Months
App Fig 7h. Death Within 12 Months
-.06 -.03 0 .03 .06
-12 -9 -6 -3 0 3 6 9 12
Month Relative to Jan 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by period


Percent Change in Chemotherapy Service Count in High Concentration Counties Relative to the January 2005 ASP Implementation Date
App Fig 8a. All Settings
App Fig 8b. Excluding Carboplatin
App Fig 8c. Physician's Office
App Fig 8d. Outpatient
-.2
-.1
0
.1
.2
-12
-9
-6
-3
0
3
6
9
12
Month Relative to January 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by setting


Percent Change in Chemotherapy Service Count in Low Concentration Counties Relative to the January 2005 ASP Implementation Date
App Fig 8e. All Settings
App Fig 8f. Excluding Carboplatin
App Fig 8g. Physician's Office
App Fig 8h. Outpatient
-.2
-.1
0
.1
-12
-9
-6
-3
0
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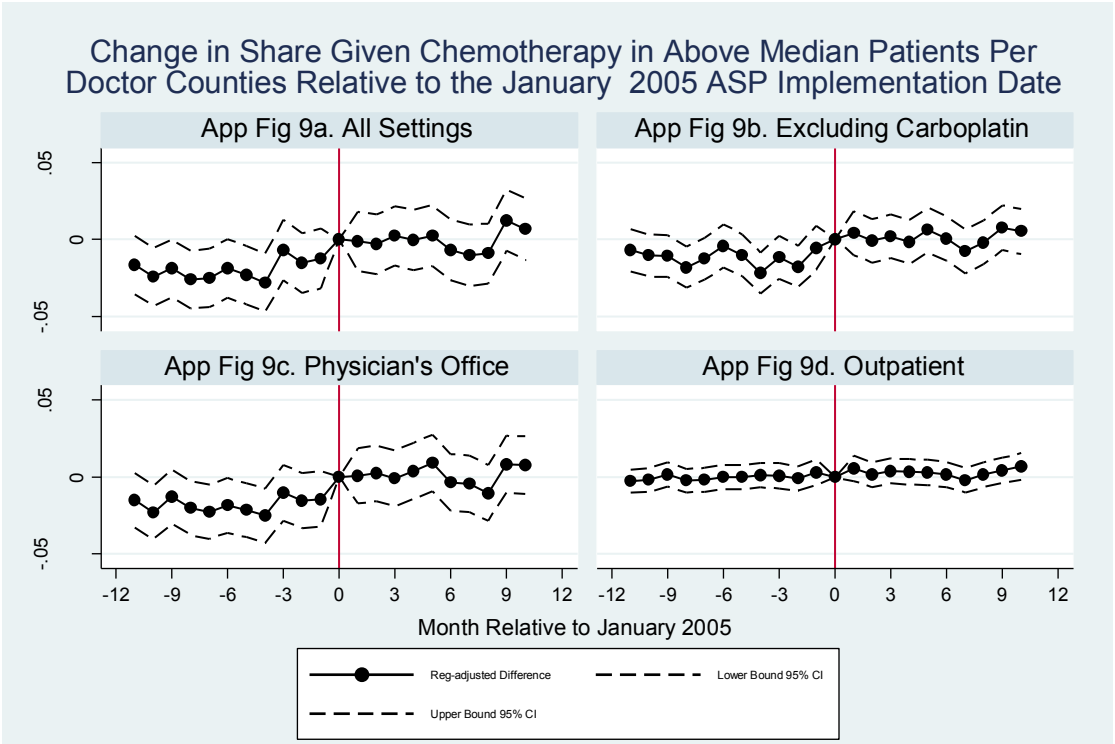
6
9
12
Month Relative to January 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by setting


Change in Share Given Chemotherapy in Above Median Patients Per Doctor Counties Relative to the January 2005 ASP Implementation Date
App Fig 9a. All Settings
App Fig 9b. Excluding Carboplatin
App Fig 9c. Physician's Office
App Fig 9d. Outpatient
-.05
0
.05
-12
-9
-6
-3
0
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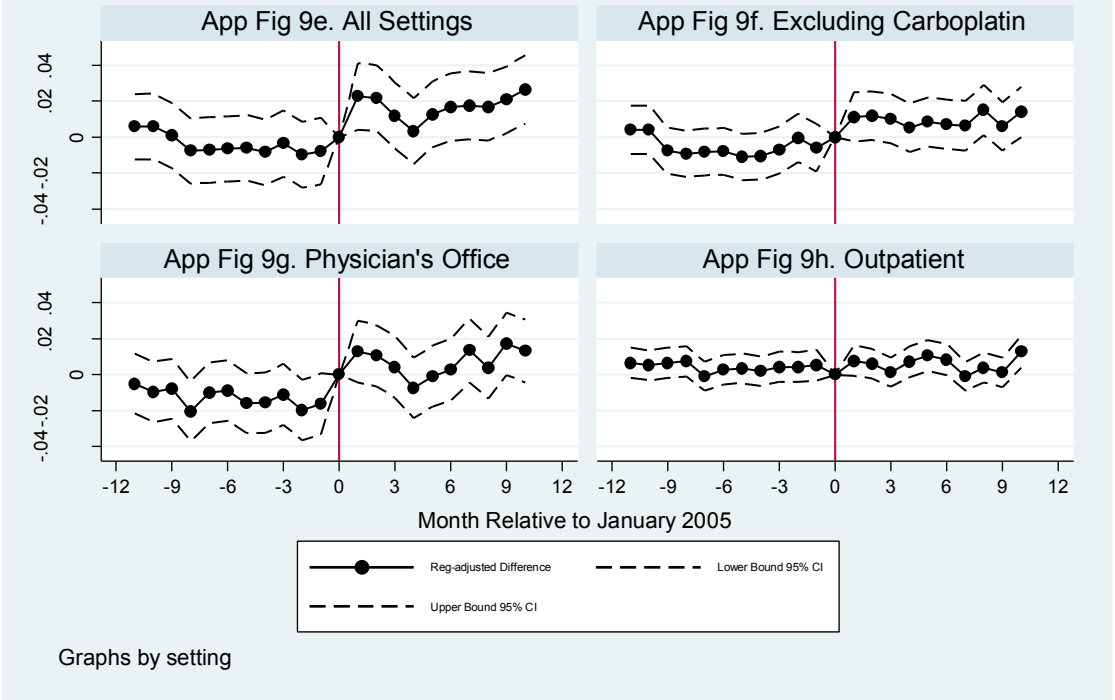
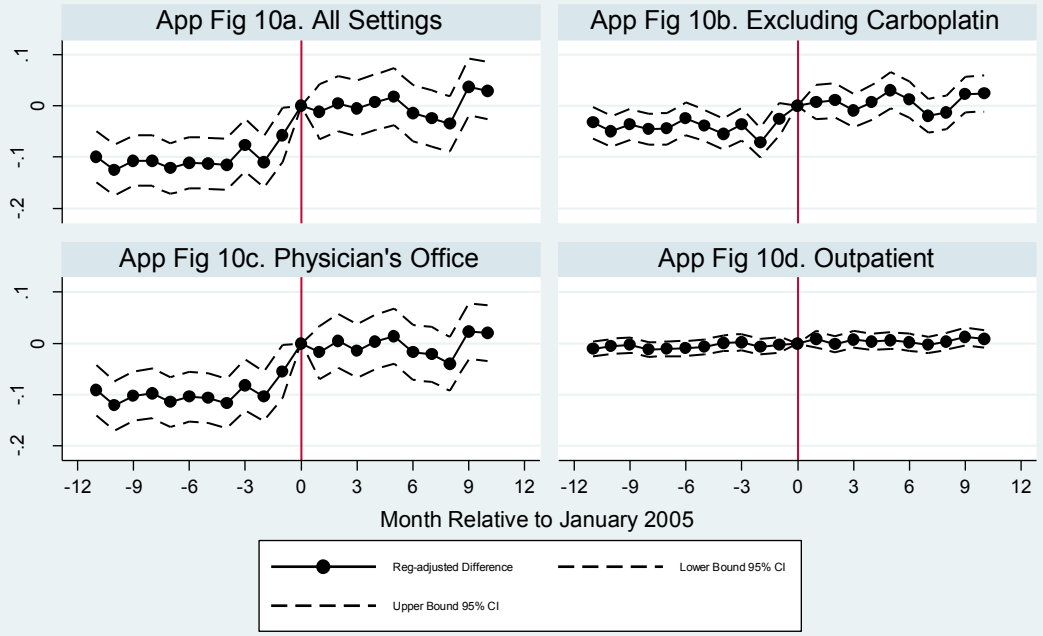
6
9
12
Month Relative to January 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by setting

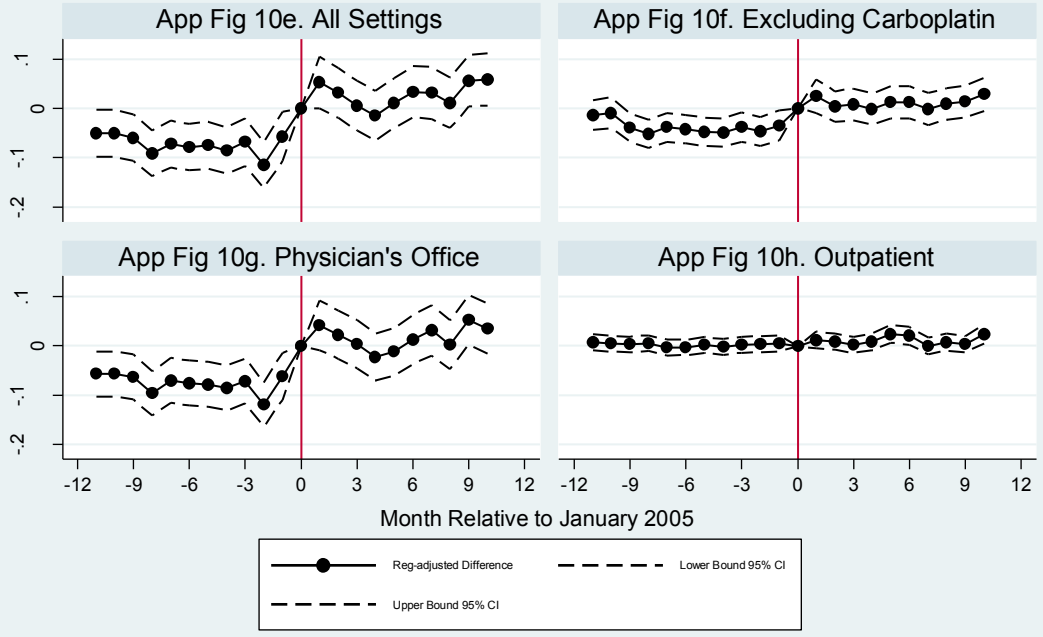

Change in Share Given Chemotherapy in Below Median Patients Per Doctor Counties Relative to the January 2005 ASP Implementation Date
App Fig 9e. All Settings
App Fig 9f. Excluding Carboplatin
App Fig 9g. Physician's Office
App Fig 9h. Outpatient
-.04
-.02
0
.02
.04
-12
-9
-6
-3
0
3
6
9
12
Month Relative to January 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by setting

Percent Change in Chemotherapy Service Count in Above Median Patients Per Doctor Counties Relative to the January 2005 ASP Implementation Date

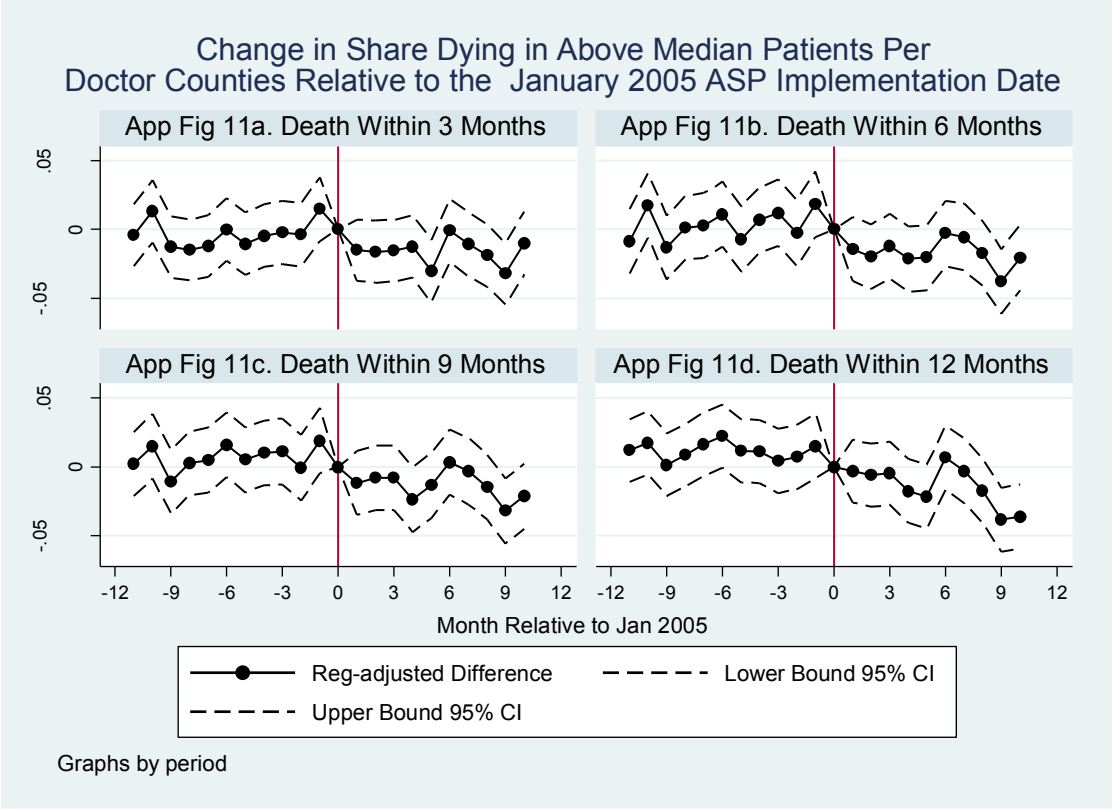


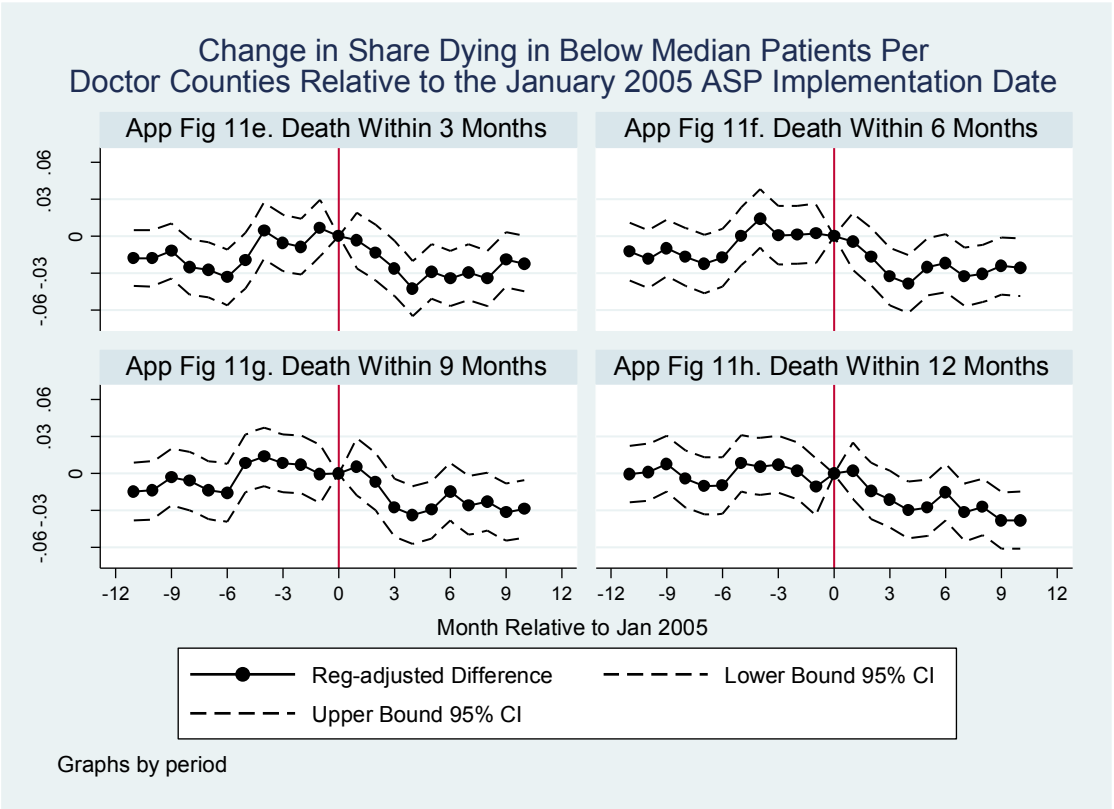
Graphs by setting

Percent Change in Chemotherapy Service Count in Below Median Patients Per Doctor Counties Relative to the January 2005 ASP Implementation Date



Graphs by setting


Change in Share Dying in Above Median Patients Per
Doctor Counties Relative to the January 2005 ASP Implementation Date
App Fig 11a. Death Within 3 Months
App Fig 11b. Death Within 6 Months
App Fig 11c. Death Within 9 Months
App Fig 11d. Death Within 12 Months
.05
0
-.05
-12 -9 -6 -3 0 3 6 9 12
Month Relative to Jan 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by period


Change in Share Dying in Below Median Patients Per
Doctor Counties Relative to the January 2005 ASP Implementation Date
App Fig 11e. Death Within 3 Months
App Fig 11f. Death Within 6 Months
App Fig 11g. Death Within 9 Months
App Fig 11h. Death Within 12 Months
.06
.03
0
-.03
-.06
-12 -9 -6 -3 0 3 6 9 12
Month Relative to Jan 2005
Reg-adjusted Difference
Lower Bound 95% CI
Upper Bound 95% CI
Graphs by period

Appendix Table1. Discontinuity in Observable Characteristics of Patients, Pre and Post Reform

	Mean Before Jan/05	Feb/04 –Nov/05	Apr/04 –Sept/05
Log Age at Diagnosis	74.1	0.0096 (0.0008)	0.010 (0.001)
Share Male	0.516	$-2.76X10^{-18}$ $(5.47X10^{-17})$	$-7.78X10^{-18}$ $(3.56X10^{-17})$
Distribution by Race/ethnicity White	0.875	-0.0030 (0.0032)	-0.0010 (0.0024)
African-American	0.091	0.0026 (0.0031)	0.0013 (0.0023)
Asian	0.0108	0.0010 (0.0062)	0.0003 (0.0007)
Other/Unknown	0.009	-.0007 (0.0003)	-.00082 (0.00046)
Hispanic	0.0103	0.0002 (0.0004)	0.0008 (0.0005)
Native American	0.0038	-0.0002 (0.0004)	-0.00082 (0.00034)
Metastatic at Diagnosis	0.289	$-9.86X10^{-18}$ $(2.09X10^{-17})$	$1.12X10^{-18}$ $(3.04X10^{-17})$
Deyo-Charlson Score	1.06	-0.006 (0.010)	-0.016 (0.013)
Number of Observations (weeks)	48	96	72

Notes: Means in column 1 are for the dependent variable for the period before ASP implementation. All cells in columns 2 - 3 present estimates from separate time-series regressions. Estimates are the coefficients on an indicator for the period that the ASP payment scheme was in effect. Regressions control for other mean patient characteristics (see Table 1), i.e. except the one used as the dependent variable or any race variables when any race category is a dependent variable, and relative-week trends that are allowed to differ on either side of the payment change. Newey-West standard errors allowing for autocorrelation up to 52-week lags are given in parentheses.

Appendix Table 2. Specification Checks of Reform-related changes in Chemotherapy Treatment within 1 months of Diagnosis

	Relative month trends		Calendar month fixed effects	
	Feb/04 –Nov/05	Apr/04 –Sept/05	Feb/04 –Nov/05	Apr/04 –Sept/05
within 1 month of Diagnosis	0.0150	0.0134	0.0175	.0183
	(0.003)	(0.002)	(0.001)	(0.002)
excluding carboplatin	0.0140	.0121	.0126	.0126
	(0.002)	(0.002)	(0.001)	(0.001)
in a Physician's Office	0.0166	0.0154	0.0181	0.017
	(0.003)	(0.002)	(0.001)	(0.002)
in an outpatient hospital clinic	-0.001	-0.001	0.002	0.002
	(0.001)	(0.001)	(0.001)	(0.001)
Number of Observations (weeks)	96	72	96	72

Notes: Each cell is based on a separate regression. Estimates are the coefficients on an indicator for the period that the ASP payment scheme was in effect. Regressions control for other mean patient characteristics (see Table 1). Columns (1) and (2) include relative-month trends that are allowed to differ on either side of the payment change and columns (3) and (4) include calendar month fixed effects. Newey-West standard errors allowing for autocorrelation up to 52-week lags are given in parentheses.

Appendix Table 3. Specification Check of Percent Change in Chemotherapy Treatment Service Counts Within 1 Month of Diagnosis After the January 2005 Payment Change

	Mean	Relative month trends		Calendar month fixed effects	
within 1 month	1.00	0.342	0328	0.445	.439
		(0.047)	(0.046)	(0.009)	(0.013)
excluding carboplatin	0.304	0.206	0.208	0238	0.214
		(0.024)	(0.022)	(0.006)	(0.007)
in a Physician's Office	0.916	0.350	0.338	0.445	0.435
		(0.048)	(0.048)	(0.009	(0.014)
in an outpatient hospital clinic	0.087	0.013	0.012	0.036	0.040
		(0.003)	(0.005)	(0.002)	(0.002)
Number of Obs (weeks)		96	72	96	72

Notes: Each cell is based on a separate regression. Estimates are the coefficients on an indicator for the period that the ASP payment scheme was in effect. Regressions control for other mean patient characteristics (see Table 1). Means in column (1) are for the dependent variable for the period before ASP implementation. Columns (2) and (3) include relative-month trends that are allowed to differ on either side of the payment change and columns (4) and (5) include calendar month fixed effects. Newey-West standard errors allowing for autocorrelation up to 52-week lags are given in parentheses.

Appendix Table 4. Changes in Chemotherapy Treatment Within 3 Months of Diagnosis After the January 2005 Payment Change: February 2004- November 2005

	Mean	Relative week trends		Relative month trends		Calendar month fixed effects	
within 3 months	0.263	0.007	0.029	0.007	0.029	0.031	.035
		(0.0037)	(0.003)	(0.0040)	(0.003)	(0.001)	(0.001)
excluding carboplatin	0.093	0.015	0.023	0.015	0.023	0.019	0.020
		(0.002)	(0.003)	(0.003)	(0.003)	(0.001)	(0.001)
in a Physician's Office	0.212	0.014	0.031	0.014	0.031	0.031	0.035
		(0.003)	(0.003)	(0.004)	(0.003)	(0.001)	(0.001)
in an outpatient hospital clinic	0.066	-0.007	-0.0004	-0.007	-0.0004	0.004	0.004
		(0.002)	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)
Quarter prior to reform indicator		N	Y	N	Y	N	Y
Number of Obs (weeks)		96	96	96	96	96	96

Notes: Means in column 1 are for the dependent variable for the period before ASP implementation. All cells in columns 2 - 7 present estimates from separate time-series regressions. Estimates are the coefficients on an indicator for the period that the ASP payment scheme was in effect. Regressions control for mean patient characteristics (see Table 1). Columns (2) and (3) include relative-week trends and (4) and (5) relative-month trends that are allowed to differ on either side of the payment change. Columns (6) and (7) include calendar month fixed effects. Newey-West standard errors allowing for autocorrelation up to 52-week lags are given in parentheses.

Appendix Table 5. Specification Checks of Reform-related changes in Chemotherapy Types within 1 months of Diagnosis: February 2004- November 2005

	Relative week trends		Relative month trends		Calendar month fixed effects	
Share of Chemotherapy Treated Patients Receiving…						
carboplatin	-0.039	-0.047	-0.038	-0.046	-0.019	-0.029
	(0.008)	(0.012)	(0.008)	(0.013)	(0.005)	(0.005)
paclitaxel	-0.074	-0.084	-0.073	-0.083	-0.059	-0.068
	(0.007)	(0.010)	(0.008)	(0.010)	(0.003)	(0.002)
docetaxel	0.026	0.036	0.026	0.036	0.014	0.013
	(0.004)	(0.005)	(0.004)	(0.005)	(0.002)	(0.003)
etoposide	-0.005	-0.025	-0.003	-0.023	-0.004	-0.007
	(0.010)	(0.007)	(0.010)	(0.007)	(0.004)	(0.003)
gemcitabine	0.007	0.007	0.007	0.008	-0.019	-0.014
	(0.003)	(0.003)	(0.003)	(0.004)	(0.002)	(0.002)
Quarter prior to reform indicator	N	Y	N	Y	N	Y
Number of Obs (weeks)	96	96	96	96	96	96

Notes: Each cell is based on a separate regression. Estimates are the coefficients on an indicator for the period that the ASP payment scheme was in effect. Regressions control for other mean patient characteristics (see Table 1). Columns (1) and (2) include relative-week trends and (3) and (4) relative-month trends that are allowed to differ on either side of the payment change. Columns (5) and (6) include calendar month fixed effects. Odd columns also include an indicator for the 3 months leading up to the reform. Newey-West standard errors allowing for autocorrelation up to 52-week lags are given in parentheses.

Appendix Table 6. Specification Checks of Reform-related changes in the Likelihood of Death

	Relative month trends		Calendar month fixed effects	
Share Dying within…	Feb/04 –Nov/05	Apr/04 –Sept/05	Feb/04 –Nov/05	Apr/04 –Sept/05
3 months of diagnosis	-0.009	-0.010	-0.012	-0.011
	(0.004)	(0.005)	(0.002)	(0.002)
6 months of diagnosis	-0.015	-0.015	-0.020	-0.020
	(0.002)	(0.003)	(0.002)	(0.003)
9 months of diagnosis	-0.011	-0.010	-0.021	-0.022
	(0.00	(0.003)	(0.003)	(0.003)
1 year of diagnosis	-0.0005	-0.0001	-0.026	-0.025
	(0.003)	(0.003)	(0.002)	(0.002)
Number of Observations (weeks)	96	72	96	72

Notes: Each cell is based on a separate regression. Estimates are the coefficients on an indicator for the period that the ASP payment scheme was in effect. Regressions control for other mean patient characteristics (see Table 1). Columns (1) and (2) include relative-month trends that are allowed to differ on either side of the payment change and columns (3) and (4) include calendar month fixed effects. Newey-West standard errors allowing for autocorrelation up to 52-week lags are given in parentheses.

Appendix Table 7. Change in Treatment and Survival: States with Above vs. Below Pre-reform Rates of Paclitaxel Use Among Patients Receiving Any Chemotherapy Treatment

	Above Median Rates of Paclitaxel Use			Below Median Rates of Paclitaxel Use		
	Mean Pre-Jan 2005	Feb/04 –Nov/05	Apr/04 –Sept/05	Mean Pre-Jan 2005	Feb/04 –Nov/05	Apr/04 –Sept/05
Share of Patients Receiving Chemotherapy Treatment within…						
1 month of diagnosis	0.177	0.018 (0.003)	0.024 (0.003)	0.154	0.014 (0.005)	0.005 (0.002)
1 month, excluding carboplatin	0.076	0.016 (0.002)	0.017 (0.003)	0.071	0.015 (0.003)	0.008 (0.003)
1 month, in a Physician's Office	0.142	0.018 (0.003)	0.024 (0.003)	0.117	0.019 (0.004)	0.014 (0.002)
1 month, in outpatient hospital clinic	0.026	-0.001 (0.001)	0.001 (0.002)	0.029	0.0001 (0.002)	0.0004 (0.001)
Share Dying Within…						
3 months of diagnosis	0.341	-0.005 (0.006)	-0.014 (0.004)	0.337	-0.001 (0.004)	-0.002 (0.003)
6 months of diagnosis	0.473	-0.021 (0.005)	-0.028 (0.007)	0.464	-0.005 (0.008)	-0.006 (0.006)
9 months of diagnosis	0.533	-0.019 (0.004)	-0.026 (0.004)	0.521	0.002 (0.007)	0.003 (0.006)
1 year of diagnosis	0. 625	-0.008 (0.003)	-0.013 (0.004)	0.612	0.012 (0.007)	0.006 (0.004)
Number of Obs (weeks	48	96	72	48	96	72

Notes: Means in column 1 and 3 are for the dependent variable for patients in states with above versus below the median rates of relative paclitaxel use in the pre-reform period. Estimates in columns 2 and 4 are from separate time-series regressions for patients from each group and are the coefficients on an indicator for the period that the ASP payment scheme was in effect. Regressions control for mean patient characteristics and relative-week trends that are allowed to differ on either side of the payment change. Newey-West standard errors allowing for autocorrelation up to 52-week lags are given in parentheses.

Appendix Table 8. Change in Treatment and Survival: Above vs. Below Median Age Patients

	Above Median Age Patients			Below Median Age Patients		
	Mean Pre-Jan 2005	Feb/04 –Nov/05	Apr/04 –Sept/05	Mean Pre-Jan 2005	Feb/04 –Nov/05	Apr/04 –Sept/05
Share of Patients Receiving Chemotherapy Treatment within…						
1 month of diagnosis	0.119	0.014 (0.002)	0.014 (0.003)	0.210	0.024 (0.003)	0.020 (.003)
1 month, excluding carboplatin	0.054	0.013 (0.002)	0.011 (0.002)	0.014	0.016 (0.002)	0.013 (0.001)
1 month, in a Physician's Office	0.094	0.014 (0.002)	0.016 (0.002)	0.164	0.024 (0.004)	0.023 (0.004)
1 month, in outpatient hospital	0.018	-0.001 (0.001)	-0.001 (0.001)	0.036	0.003 (0.001)	0.004 (0.002)
Observations (weeks)	48	96	72	48	96	72

Notes: Means in column 1 and 4 are for the dependent variable for patients above versus below the median age patient at diagnosis in the period prior to the reform. see notes to Tables 2-4. See notes to Tables 2-4 for more details.

Appendix Table 9. Change in Treatment and Survival: Counties with a High vs. Low Concentration of Providers Administering Chemotherapy Treatment

	Counties with High Provider Concentration			Counties with Low Provider Concentration		
	Mean	Feb/04 –Nov/05	Apr/04 –Sept/05	Mean	Feb/04 –Nov/05	Apr/04 –Sept/05
Share of Patients Receiving Chemotherapy Treatment within…						
1 month of diagnosis	0.169	-0.0004 (0.005)	0.001 (0.007)	0.165	0.020 (0.002)	0.019 (0.003)
1 month, excluding carboplatin	0.076	0.004 (0.003)	0.002 (0.005)	0.073	0.016 (0.001)	0.014 (0.002)
1 month, in a Physician's Office	0.122	0.008 (0.005)	0.010 (0.007)	0.132	0.019 (0.003)	0.019 (0.003)
1 month, in outpatient hospital clinic	0.045	-0.010 (0.002)	-0.009 (0.002)	0.024	0.003 (0.001)	0.004 (0.001)
Percent Change in Chemotherapy Service Counts within…						
1 month of diagnosis	0.973	0.203 (0.045)	0.185 (0.045)	1.01	0.389 (0.031)	0390 (0.030)
1 month, excluding carboplatin	0.320	0.089 (0.038)	0.085 (0.026)	0.301	0.232 (0.021)	0.240 (0.014)
1 month, in a Physician's Office	0.821	0.242 (0.045)	0.231 (0.044)	0.933	0.389 (0.034)	0.392 (0.033)
1 month, in outpatient hospital clinic	0.152	-0.043 (0.024)	-0.049 (0.017)	0.075	0.029 (0.005)	0.028 (0.008)
Observations (weeks)	48	96	72	48	96	72

Notes: Means in column 1 and 3 are for the dependent variable for patients in counties with low vs. high concentration. The mean for service counts is without logs. High concentration is defined as an HHI greater than 0.25. Estimates in columns 2 and 4 are from separate time-series regressions for patients from each group and are the coefficients on an indicator for the period that the ASP payment scheme was in effect. Regressions control for mean patient characteristics and relative-week trends that are allowed to differ on either side of the payment change. Newey-West standard errors allowing for autocorrelation up to 52-week lags are given in parentheses

Appendix Table 10. Change in Treatment and Survival: Counties with Above vs. Below Median Numbers of Patients Per Provider Administering Chemotherapy Treatment

	Above Median Patients Per Provider			Below Median Patients Per Provider		
	Mean Pre-Jan 2005	Feb/04 –Nov/05	Apr/04 –Sept/05	Mean Pre-Jan 2005	Feb/04 –Nov/05	Apr/04 –Sept/05
Share of Patients Receiving Chemotherapy Treatment within…						
1 month of diagnosis	0.173	0.013 (0.004)	0.013 (0.003)	0.159	0.019 (0.001)	0.017 (0.003)
1 month, excluding carboplatin	0.076	0.015 (0.002)	0.014 (0.003)	0.071	0.014 (0.003)	0.010 (0.003)
1 month, in a Physician's Office	0.138	0.017 (0.002)	0.018 (0.004)	0.122	0.019 (0.003)	0.019 (0.003)
1 month, in outpatient hospital clinic	0.025	-0.0004 (0.001)	-0.0004 (0.001)	0.030	-0.0001 (0.001)	0.001 (0.002)
Percent Change in Chemotherapy Service Counts within…						
1 month of diagnosis	1.06	0.336 (0.048)	0.328 (0.061)	0.942	0.400 (0.035)	0383 (0.036)
1 month, excluding carboplatin	0.322	0.185 (0.030)	0.203 (0.029)	0.286	0.258 (0.022)	0.245 (0.017)
1 month, in a Physician's Office	0.987	0.339 (0.043)	0.337 (0.062)	0.844	0.411 (0.036)	0.395 (0.037)
1 month, in outpatient hospital clinic	0.077	0.019 (0.008)	0.012 (0.007)	0.097	0.013 (0.004)	0.015 (0.007)
Share Dying Within…						
3 months of diagnosis	0.353	-0.009 (0.006)	-0.019 (0.004)	0.335	-0.002 (0.005)	0.008 (0.005)

6 months of diagnosis	0.474	-0.015 (0.006)	-0.022 (0.006)	0.463	-0.009 (0.004)	-0.003 (0.004)
9 months of diagnosis	0.532	-0.008 (0.006)	-0.028 (0.003)	0.522	-0.004 (0.003)	0.001 (0.005)
1 year of diagnosis	0. 635	-0.008 (0.008)	0.024 (0.003)	0.616	0.008 (0.004)	0.012 (0.004)
Number of Obs (weeks)	48	96	72	48	96	72

Notes: Means in column 1 and 3 are for the dependent variable for patients in counties with above vs. below median number of patients per provider. Estimates in columns 2 and 4 are from separate time-series regressions for patients from each group and are the coefficients on an indicator for the period that the ASP payment scheme was in effect. Regressions control for mean patient characteristics and relative-week trends that are allowed to differ on either side of the payment change. Newey-West standard errors allowing for autocorrelation up to 52-week lags are given in parentheses