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How Does Competition Impact the Quality and Price of Outpatient Service Facilities?
A Case Study of the U.S. Dialysis Industry

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Abstract

We examine whether and to what extent consolidation in the largely for-profit U.S. dialysis industry has affected patient outcomes, clinical practices, and prices charged to the privately-insured. We make use of detailed facility data for the period 2000-2009, during which the market share of the two industry leaders increased from just over one-third to nearly two-thirds. We exploit the differential impact of two large national acquisitions on local market concentration to estimate the causal effect of concentration on a broad set of measures, including mortality rates, dialysis adequacy, staffing ratios, and price per treatment. We find no statistically or economically significant effects of competition on any outcome or practice measure. Preliminary analysis suggests consolidation may have led to higher private prices.

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I. Introduction

The U.S. healthcare system is designed around the premise and promise of robust competition throughout the value chain of production. Even as doubts accumulate about the ability of market-based healthcare systems to deliver superior quality at lower cost, other countries are introducing competition-based reforms in an attempt to harness the benefits arising from the pursuit of rational self-interest. Among the leading examples are "Choose and Book," an initiative to expand patients' hospital choices in the UK, and the 2006 decision by the Netherlands to mandate private health insurance coverage.

Of course, whether competition "works" depends on the particulars of the environment, as well as on the outcomes defining success. Most prior research focuses on the U.S. hospital industry, in particular on the effect of competition on prices negotiated with private insurers. This research has yielded conclusive evidence that competition restrains prices.[1] The effects of competition on the quality of inpatient care are less conclusive, in part because available measures of quality are crude (e.g. inpatient mortality) or have ambiguous interpretations (e.g. length of stay, number of procedures), and in part because it is difficult to separately identify quality from price effects. In this paper, we revisit the question of whether competition enhances quality of care in a setting with uniquely comprehensive quality data, and prices largely fixed by Medicare: freestanding kidney dialysis clinics.

There are approximately 5,760 dialysis facilities in the U.S. today.[2] These outpatient facilities – over 80 percent of which are for-profit - provide frequent dialysis treatments to over 400,000 Americans suffering from end-stage renal disease (ESRD), an affliction characterized by severely diminished functioning of the kidneys.[3] To survive, patients with ESRD must undergo kidney transplantation or submit to frequent dialysis, a procedure which mimics the role of the kidneys by eliminating waste from the patient's blood. The Medicare program, which

[1] Gaynor and Town (2011) lay out the theoretical foundations of this literature, and provide a thorough and critical summary of the associated empirical work.

[2] All figures are for 2008. Source: 2011 USRDS Annual Data Report, Volume 2, available at http://www.usrds.org/2011/pdf/v2_00_intro_11.pdf: Appendix Table 11.2, Appendix Table 10.19, Table 11.3, Table 11.5 and authors' calculations.

[3] The number of ESRD patients is growing by 3 percent per year (*ibid)*. By comparison, approximately 622,000 percutaneous transluminal coronary angioplasties (PTCAs) are performed in the U.S. each year, and the growth rate is negative ("Heart Disease and Stroke Statistics: 2011 Update: A Report from the American Heart Association," *Circulation* 2011;123:e18-e209. Downloadable at http://circ.ahajournals.org/content/123/4/e18.).

extended coverage to ESRD patients (regardless of age) in 1972, foots the bill for more than 80 percent of dialysis patients.[4] Just 1.3 percent of Medicare enrollees have ESRD, but they account for 6.2 percent of total Medicare spending (including Part D). Slightly more than half of this spending is attributable to dialysis.

While the incidence of ESRD has increased fairly steadily since the 1980s – when the data was first systematically gathered by the government-funded U.S. Renal Data System (USRDS) – the number of dialysis facilities began skyrocketing in 1990. This surge closely followed the introduction of "erythropoiesis stimulating agents" (ESAs), injectable drugs for combating anemia.[5] ESAs dramatically boosted patients' quality of life, and also proved extremely lucrative for facilities to administer, as it was reimbursed on a generous fee-for-service basis. Utilization of ESAs is particularly high among the largest for-profit chains (Thamer et al. 2007). In 2007, the *New York Times* reported that DaVita Inc., one of the two leading chains, earned 25 percent of its revenues and 40 percent of its profits through this drug alone.

In addition to rapid industry growth, the dialysis sector has undergone a massive wave of consolidation in recent years, as illustrated in **Figure 1.** The top two firms in 1996, Fresenius and Vivra Renal Care, accounted for 20 percent and 8 percent of approximately 3,000 facilities, respectively. The national four-firm concentration ratio stood at 35 percent of facilities.[6] By 2009, the top two firms jointly accounted for nearly two-thirds of the market (32 percent for Fresenius, 30 percent for DaVita), notwithstanding a near-doubling in the total number of facilities.[7] These figures are slightly higher today, as Fresenius and DaVita have since acquired the third and fourth-largest chains, respectively.[8]

[4] To qualify, individuals must be eligible to receive benefits from Social Security, the Railroad Retirement Board, or as a "Medicare-qualified government employee," or be the spouse or dependent child of someone who qualifies. According to the USRDS, in 2008 83 percent of ESRD patients were enrolled in Medicare (*ibid).*

[5] The vast majority of dialysis treatments take place in dialysis facilities. Less than 8 percent of patients dialyze at home (7 percent using peritoneal, and <1 percent using hemodialysis modalities). (USRDS Annual Data Report, Volume 2, Chapter 4 http://www.usrds.org/2010/view/v2_04.asp). home hemodialysis population reached 3,826.

[6] 1998 USRDS Annual Data Report, Chapter 11 and author's calculations from Cost Reports (http://www.usrds.org/download/1998/ch11.pdf). The 3rd and 4th largest chains were DaVita's predecessor (4%) and Dialysis Clinic Inc (3.7%)

[7] 2011 USRDS Annual Data Report, Volume 2 2011, Appendix Table 10.1.

[8] Fresenius acquired Liberty Dialysis, a 260-clinic chain headquartered in Mercer Island, WA, while DaVita acquired DSI, a 106-clinic chain based in Nashville, TN. To preserve competition in local geographic markets, the Federal Trade Commission required divestitures in both mergers. For details, see http://www.ftc.gov/opa/2012/02/fresenius.shtm and http://www.ftc.gov/opa/2011/09/DaVita.shtm.

The recent consolidation in the dialysis sector, coupled with fixed prices for most patients, presents an excellent setting to explore how competition affects quality.[9] In addition, the quality measures available for kidney dialysis are unprecedented in terms of their detail and medical relevance. Detailed "Dialysis Facility Reports" (DFRs) have been prepared by the Kidney Epidemiology and Cost Center at the University of Michigan and provided to facilities since 1995.[10] For the past several years, only a small (and censored) subset of statistics has been publicly available through Medicare's online "Dialysis Compare" tool. However, in 2010 an investigative journalist employed by the nonprofit organization *ProPublica* filed a successful request under the Freedom of Information Act to obtain the original, detailed DFRs. This journalist, Robin Fields, provided us access to the DFRs for the period 2002-2010 (spanning data years 1998-2009), which we assembled into a comprehensive database.

To evaluate the effects of competition on a variety of outcome measures, we make use of plausibly exogenous changes in local geographic market structure generated by two large, national acquisitions in 2005-2006, one each by market leaders Fresenius and DaVita. The impact of these acquisitions on market structure (as measured by the Hirschman-Herfindahl index, or HHI) varied with the pre-acquisition market shares of the targets and acquirers. The large and national nature of the transactions mitigates concerns that consolidation within any particular market was driven by unobservable factors, and it also permits us to separate the impact of ownership changes from that of local market structure. While our focus is on the effects of competition on a range of dialysis quality measures, we also pursue a secondary analysis of the effects of competition on private dialysis prices. The privately-insured account for less than 20 percent of patients, but they generate a disproportionate share of revenues (e.g., 35 percent of DaVita's revenues in 2009).[11]

We find that the 2005-2006 mergers resulted in sizeable increases in HHI for dozens of markets. Specifically, among the 91 Hospital Service Areas (as defined by the Dartmouth Atlas) in which at least one of the targets and its acquirer overlapped, the median predicted increase in HHI was 1000 points (which, according to our estimates, translated into a post-merger increase

[9] The nursing home industry is another such example, and is the subject of related empirical IO work by Lu (2012) and Chen (2009).
[10] The Dialysis Facility Reports have been funded directly by the Centers for Medicare and Medicaid Services since 1999. For details, see http://www.sph.umich.edu/kecc/assets/documents/HistoryofKECC.pdf.
[11] 2009 DaVita Annual Report (http://phx.corporate-ir.net/External.File?item=UGFyZW50SUQ9Mzg3MzkwfENoaWxkSUQ9Mzg5ODU0fFR5cGU9MQ==&t=1)

of about 880 points). Notwithstanding these large increases in HHI - in markets which were fairly concentrated to begin with - we find fairly precise "zero effects" of changes in HHI on a broad set of outcome measures.

The paper proceeds as follows. Section II summarizes prior research and provides background information on dialysis and the industry that surrounds it. Section III describes the data. We examine the evolution of local dialysis market concentration in Section IV. Section V presents our analyses of the relationship between changes in local market concentration and dialysis quality. We discuss results for fixed-effects models as well as reduced-form models which make use of the variation in concentration induced by the two national acquisitions of 2005-2006. Section VI describes our analyses of the relationship between changes in concentration and private dialysis prices, which are determined via negotiation between private insurers and dialysis providers. Section VII concludes.

II. Background and Prior Research

A. What is Dialysis and How is Dialysis Quality Measured?

There are many causes of ESRD, but only one cure: kidney transplant. However, less than 20 percent of those diagnosed with ESRD receive a transplant, owing to organ shortages as well as comorbidities that render the surgery too risky.[12] Kidneys perform two main functions. First, they filter the blood, producing urine as a waste product while maintaining correct blood Ph and blood pressure. Second, kidneys secrete a variety of hormones, including erythropoietin, which stimulates the production of red blood cells, and calcitrol, an essential form of vitamin D.

A basic dialysis treatment mimics the first function of the kidneys, either by passing the blood through a machine and returning it to the patient (hemodialysis) or by filtering it inside the patient's abdominal cavity, which can be accomplished by inserting dialysis solution into the cavity and draining it 4-6 hours later (peritoneal dialysis). Roughly 90 percent of patients choose hemodialysis as their treatment modality; of these, 99% rely on in-center hemodialysis.[13]

[12] There were 17,736 kidney transplants performed in the US in 2009 (2011 USRDS Annual Data Report, Volume 2, Chapter 7). By contrast, over 110,000 individuals were diagnosed with ESRD. Only 28,494 of these were added to a transplant waiting list.

[13] 2008 USRDS Annual Data Report, Volume 2, Chapter 4 (http://www.usrds.org/2008/view/esrd_04.asp)

The second function of the kidneys is fulfilled by a cocktail of drugs, many of which are injected during dialysis treatments. One category in particular has attracted a great deal of attention: synthetically produced erythropoeisis-stimulating agents (alternatively known as ESAs, erythropoietin, epoetin, epogen or EPO). Following the 1989 FDA approval of epoetin, the popularity of dialysis surged. In 2009, epoetin was Medicare's largest drug expenditure, with spending of ~$2.71 billion, nearly 70 percent of which was incurred by ESRD patients.[14] Epoetin has dramatically reduced anemia in dialysis patients, but its overuse is linked to increases in adverse cardiovascular events (stroke, heart attack, and heart failure).[15] In the following section, we summarize the research on usage of epoetin across facilities of different types.

Given that urea removal is a main objective of dialysis, a common measure for dialysis efficacy is the *urea reduction ratio* (URR), the fraction of urea in a patient's blood that is eliminated during a given dialysis treatment. Anemia, or lack of red blood cells, is measured by levels of *hematocrit,* which is the fraction of the blood composed of red blood cells, or *hemoglobin,* which is the iron-containing pigment in red blood cells.[16] Given the high mortality rate of dialysis patients – around 20 percent over the course of a year of treatment – mortality rates are also indicators of dialysis facility performance.

Although the U.S. spends more per ESRD patient than any other country (Dor et al. 2007), mortality rates of U.S. ESRD patients are relatively high. For example, Goodkin et al. (2003) report that the relative risk of death, adjusted for demographics and 15 classes of comorbidities, is 3.78 for the U.S., versus Japan's 1.0 and Europe's 1.33. The unadjusted relative risks of death stand at 5.34 to 1 for the U.S. versus Japan (and 3.12 to 1 for Europe versus Japan). Robinson and Port (2009) review the literature on these international differences. They conclude that the figures cited above suffer from important shortcomings in risk-adjustment. For example, U.S. patients receive kidney transplants at a much higher rate than Japanese patients, leaving behind a comparatively sicker ESRD population (because transplant candidates are

[14] This combines expenditure on two types of epoetin: alpha and beta. ESRD epoetin spending in 2009 was $1.87 billion according to the 2011 USRDS Annual Data Report, Volume 2, Chapter 11, p. 282 (http://www.usrds.org/2011/pdf/v2_ch011_11.pdf). Medicare Part B epoetin spending in 2009 was $840 million according to Medpac (http://www.medpac.gov/documents/Jun11DataBookEntireReport.pdf, p157-158 and authors' calculations).

[15] These risks are prominently featured at http://www.epogen.com/.

[16] Hematocrit is a percentage, whereas hemoglobin is measured measured in grams per deciliter. The correlation is exceedingly high, with hematocrit approximately equal to 3 * hemoglobin.

relatively healthy).[17] In addition, higher "background mortality rates" in the U.S. contribute to worse outcomes for ESRD patients.[18] However, it appears that a good portion of the differences in outcomes are attributable to differences in practice patterns. For example, Pisoni et al. (2009) find that over half of the differences in risk-adjusted mortality rates between the US and Europe can be explained by differences in vascular access methods.[19] Port et al. (1998) find that changes in practice patterns contributed to substantial decreases in mortality rates in the US between 1986 and 1997. Last, there is emerging evidence that longer and more frequent dialysis sessions, which are commonplace in much of Europe, Australia, New Zealand, and Japan, lead to improved health and longevity (e.g., Tentori et al. 2012).

A limited set of quality data, along with facility descriptive information (e.g. address, hours, ownership type) has been available online through the *Dialysis Facility Compare* website (www.dialysiscompare.gov) since 2002. In 2011, *Dialysis Facility Compare* included the following three measures: (1) URR; (2) percent of Medicare patients with average hemoglobin<10 and >12 (the latter is an indicator of excess epoetin use); (3) crude categorization of *standardized mortality rates*, which are adjusted for patient demographics and comorbidities. In 2011, *standardized mortality rates* reflected experience between 2006 and 2009. Just 9 percent of facilities were marked "better than expected," and 11 percent "worse than expected," with the remainder characterized "as expected." Snyder and Ramanarayanan (2012) investigate the implications of this reporting discontinuity on clinic quality. They find that firms classified in the "worse than expected" category subsequently improve quality more than clinics narrowly missing this categorization. The evidence suggests this quality improvement is largely due to strategic patient selection. Portending our quality results, they find little evidence that consumers are responding to the quality information, suggesting quality inelastic demand and therefore potentially small quality responses to changes in the competitive environment.

As compared to what is reported online in *Dialysis Facility Compare*, the Dialysis Facility Reports are extraordinarily detailed, with some 18 pages of customized text and tables

[17] Interestingly, Wong et al. (1990) find that differences between Japanese and American dialysis mortality rates are mirrored by differences within the US in Asian versus Caucasian mortality rates.
[18] Pisoni et al. (2009) find that background mortality rates explain nearly half of the international variation in risk-adjusted mortality rates. Cardiovascular disease mortality rates are particularly highly correlated between the general and dialysis populations. Substantial variation in mortality rates remains.
[19] From "best" to "worst", vascular access methods are fistula, graft and catheter. The authors find that other practice pattern adjustments have a much smaller effect on mortality differences.

per facility. An excerpt is included as **Figure 2**.. In Section III below, we discuss the measures we utilize in our analyses.

B. Dialysis Reimbursement and Industry Overview

The federal government plays a uniquely prominent role in funding healthcare for individuals with ESRD. A 1972 law ensured that nearly anyone suffering from ESRD would gain Medicare coverage and that Medicare coverage would cover three dialysis treatments per week.[20] However, Medicare is a secondary payer during the first 30 months of eligibility for ESRD patients with private insurance.[21] Until 2011, Medicare reimbursed dialysis treatment - excluding most drugs - at a roughly constant nominal price (called a "composite rate") of ~\$130.[22] Thus, inflation-adjusted Medicare reimbursement for dialysis dropped dramatically over time, from \$670 in 1973 to \$132 in 2010, measured in constant 2010 dollars.

Prices paid by private insurers are roughly twice as high (USRDS 2010). Until 2011, injectable drugs such as epoetin and vitamin D replacements were separately reimbursable. These represented a substantial source of industry revenues and profits. For example, in 2006, 35% of DaVita's revenues were for physician-prescribed pharmaceuticals, with 71% of that revenue coming from epoetin.[23] Concern about the soaring expense, overutilization, and associated health risks led to the introduction of a "bundled" composite base rate of \$229.63 in 2011 (inclusive of drugs).[24]

Nephrologists provide medical oversight for dialysis facilities in exchange for annual compensation said to range from \$20,000 - \$200,000 annually.[25] The median reported compensation in the 2009 Medicare Cost Reports (described below) is approximately \$75,000, however experts we interviewed suggest these figures are understated. Medical directors are a

[20] 92-93 % of the US population would be eligible for Medicare ESRD coverage if they contracted ESRD (Nissenson & Rettig, 1999). Most ineligibility is because of work requirements to qualify for Medicare.
[21] The secondary payer provision first appeared (with an 18-month requirement) as part of the Omnibus Budget Reconciliation Act of 1981. This period was extended to 30 months by the Balanced Budget Act of 1997.
[22] Medicare reimbursed \$138 from 1973 until the implementation of PPS in 1983, at which point basic wage adjustments were added. Base reimbursement rates changed minimally in nominal terms from 1983 (\$123) to 2010 (\$132).
[23] Per the Annual Report of DaVita, Inc. (2006).
[24] CMS final rule, accessed at: http://edocket.access.gpo.gov/2010/pdf/2010-18466.pdf
[25] According to the 2009 Medicare Cost Reports, the median number of "hours per week needed to perform the job of medical director" is 10. Compensation range reported in: "Rivals wary of dialysis giant DaVita's aggressive business style," *Denver Post* 7/14/2009.

primary source of facility referrals. Clinics also reportedly compete for patients via amenities (e.g. personal TVs and heated massage chairs). As we discuss below, we are not aware of any published studies on whether quality of care (or amenities) affect clinic selection.

Dialysis facilities may be "freestanding" or "hospital-based." The share of freestanding facilities has increased over time, from 77 percent in 1998 to 86 percent in 2009.[26] Both facility types draw comparable patient populations, but their ownership structures differ: 91 percent of freestanding facilities are for-profit, whereas 95 percent of hospital-based facilities are not-for-profit. Relatedly, most freestanding facilities are in a chain with 100+ clinics (78% in 2009), while few hospital facilities are (under 2%).[27]

C. Prior Research

This paper draws on two distinct streams of research: the IO and health-IO literature on the effects of competition on quality, and the health-services literature on the dialysis industry. We discuss each in turn below.

C. 1 *Research on Competition in Quality in Healthcare Settings*

Most of the empirical evidence on the relationship between competition and vertical quality comes from healthcare settings.[28] In the dialysis industry, prices are fixed for the vast majority of patients, so our discussion focuses on the literature in which prices are fixed.[29] If prices are fixed and quality affects demand, then quality should increase with competition. The relevant literature is well-summarized in Gaynor and Town (2011) and Dranove (2011); here we provide a brief overview in order to provide context for our contribution.

[26] 2004 USRDS Annual Data Report, Volume 2, Table J.7 and 2011 USRDS Annual Data Report, Volume 2, Figure 10.1 This reduction is comprised of a small absolute decline in the number of hospitals with facilities, and a large increase in the number of freestanding facilities.

[27] Figures are calculated by the authors using 2009 data from the 2011 USRDS Annual Data Report, Volume 2, Figure 10.1.

[28] Outside of healthcare, there is a small empirical literature on the effects of competition on quality when price is fixed, e.g. Hoxby (2000) on schools and Mazzeo (2003) on airlines.

[29] The literature with market-determined prices is voluminous. Some studies find that competition improves quality (e.g., Sohn and Rathouz 2003), some find no effect (e.g., Ho and Hamilton, 2000; Capps, 2005) and some find that competition diminishes quality (e.g., Propper, Burgess, and Green, 2004; Burgess, Propper and Gossage, 2008). When firms choose both prices and quality, theoretical predictions about the effects of competition on quality are ambiguous.

The relevant papers can be divided into those which exploit plausibly exogenous shocks to market structure to identify the effect of competition on quality, and those which exploit plausibly exogenous shocks to providers – who are located in markets of different, but fixed, competitiveness – to identify this effect. The seminal paper in the first category is Kessler and McClellan (2000).[30] The authors estimate the impact of competition on mortality of Medicare patients admitted to the hospital following acute myocardial infarction (AMI), or heart attacks. Their estimating equation follows the traditional "Structure Conduct Performance" (SCP) framework, in which an outcome is regressed on a measure of area competition; in their case they regress individual mortality outcomes on a hospital-year HHI. They construct this HHI using *predicted* rather than actual patient flows, as the latter would lead to an endogenous estimate: if patients are willing to travel further for particularly good hospitals, better hospitals will mechanically be placed in larger markets with more competitors. The authors include hospital fixed effects in their SCP regressions, so that the effect of interest is identified by changes in competitiveness arising from entry, exit, and mergers of local hospitals, changes in the location of AMI patients, and changes in preferences regarding distance to the hospital. Kessler and McClellan find competition improves the quality of inpatient AMI care, particularly in those areas with greater penetration of managed care. The paper represents a major improvement over the previous literature, however some of the sources of identifying variation are subject to the well-known critiques of the SCP methodology (Demsetz 1973). For example, lower-quality hospitals may be more likely to exit, leading to an upward-biased coefficient between market level HHI and market level quality. A more recent literature (e.g. Town and Liu (2003) and Gaynor and Vogt (2003)) estimates the effect of competition on various outcomes by building a structural model of an industry and simulating the effects of entry or exit.

The second stream of empirical research focuses on responses of incumbents in various market structures to a common external shock. The U.K.'s "Choose and Book" policy referenced earlier has elicited a few recent papers, including Gaynor, Moreno-Serra, and Propper

[30] Research prior to Kessler and McClellan (2000) analyzed how competition affects service offerings, capacity, costs and prices, rather than how competition affects outcomes. Dranove and White (1994) summarize and critique this older literature. Gowrisankaran and Town (2003) and Kessler and Geppert (2005) rely upon a similar methodology. Gowrisankaran and Town (2003) find that competition improved quality for private insurance patients, while worsening it for Medicare patients. Kessler and Geppert (2005) study how the effects of competition on quality vary with illness severity.

(2010) and Cooper, Gibbons, Jones, and McGuire (2011). Like the U.S. Medicare program, the British National Health Service utilizes prospective payment for inpatient admissions. However, until 2006 most patients were not given a choice of inpatient hospital. The so-called "Choose and Book" reforms created tools to facilitate comparisons across hospitals and required referring physicians to give prospective hospital patients 5 choices for scheduled inpatient visits. In addition, local public agencies were permitted to engage in selective contracting with hospitals. Both Gaynor, Moreno-Serra, and Propper (2010) and Cooper et al. (2011) find that Choose and Book decreased risk-adjusted mortality rates among AMI admissions[31] to a greater extent in more competitive markets, as measured by hospital-specific HHIs.[32]

Methodologically, the papers most similar to ours are Dafny, Duggan and Ramanarayanan (2012) and Chen (2008). Dafny, Duggan and Ramanarayanan (2012) exploit a large, national merger in the health insurance industry to study to effect of consolidation on health insurance premiums. Chen (2008) uses statewide exits of a nursing home chains to study the effects of changes in local competition on staffing levels. Our paper complements these earlier projects by providing insight into the competitive landscape of a large, lucrative, and growing industry, and in so doing uncovers facts and patterns which may pertain to a broader set of industries dominated by national chains.

C.2 *Research on the U.S. Dialysis Industry*

Prior research on the U.S. dialysis industry focuses on two key areas: (1) the relationship between ownership/organizational form (i.e., for-profit status and chain size) and practice patterns/patient mortality; (2) the effects of reimbursement policies on practice patterns/patient mortality. Several studies have documented inferior quality in for-profit facilities, e.g. higher hospitalization rates (both overall, and for heart failure and volume overload; Dalrymple et al. 2013), lower transplantation rates (Garg et al, 1999), excessive epoetin doses (Thamer et al., 2007), and greater use of more expensive epoetin-delivery methods (Thamer et al., 2006). Most studies also find higher mortality rates in for-profit facilities, although the literature suffers

[31] AMI mortality is a common quality measure, and a strong indicator of overall hospital quality (Gaynor 2006). Given that AMI patients are taken to the closest suitable facility, using AMI mortality as a quality measure mitigates concerns about selection bias and demand inducement for elective procedures following the reform.

[32] Gaynor et al. instrument for hospital-specific HHI using Kessler and McClellan's methodology, however the key identifying variation comes from differences in the competitiveness of hospitals' markets prior to the reform, which in turn affect hospitals' responses to the reform.

from selection biases (Brooks et al. 2006). In addition, there is substantial variation in standardized mortality rates within chains (Zhang, Cotter, and Thamer (2011).

There are a handful of studies which explore the effects of local industry market structure on provider behavior. Held and Pauly (1983) report higher levels of measurable ammenities and higher costs in more competitive markets. These findings are corroborated by Hirth, Chernew and Orzol (1999), who also find faster adoption of quality-enhancing technology in these markets. Interestingly, adoption of cost-cutting technologies is not sensitive to market structure. Finally, Farley (1996) finds that facilities in more competitive markets are more likely to treat patients who are “clinically marginal and otherwise might not qualify for treatment.”

To our knowledge, this is the first longitudinal study exploring the impact of changes in competition on changes in dialysis quality and price. As noted above, we pursue both OLS and IV analyses to gain a more complete picture of the relationship between quality or price and market structure.

III. Data

A. Dialysis Facility Data

We assembled a database of facility-year data for the period 2000-2009 using four publicly-available sources: the Dialysis Facility Reports, the Dialysis Cost Reports (for freestanding facilities only), the Hospital Cost Reports (for hospital-based facilities only), and Dialysis Facility Compare.[33] We describe each of these sources in turn below.

The Dialysis Facility Reports (DFRs) are our primary source of data. As previously noted, the DFRs were not available to the public until 2010, when ProPublica successfully filed a request under the Freedom of Information Act and received over 40,000 PDF files of these reports, for the years 2002-2010. The DFRs contain a rich set of facility-specific information about patients, utilization, and outcomes. The data are presented alongside state, regional, and national averages to facilitate benchmarking. Each DFR reports data separately for four years, with a one-year lag (e.g., the 2010 report contains data for 2006-2009). We retained the most recent vintage for each calendar year, and merged these data (by unique provider ID) to the

[33] The data on the Dialysis Facility Reports extends back to 1998. We exclude 1998 and 1999 – the format and contents of the reports change substantially across those years. In particular, they lack information on whether different facilities are consolidated into one report, making it difficult to match facilities across the multiple datasets.

Dialysis Cost Reports and Hospital Cost reports. Our sample spans the years 2000-2009, which provides an ample pre and post-period with respect to the two consolidations of interest.[34]

The Dialysis Cost Reports include identifying information (e.g., provider ID number, location, profit status, owner identity), and detailed financial and operational information (e.g., staffing levels, patient volumes). Dialysis clinics that are sponsored by hospitals are consolidated into the hospital's Cost Report and therefore do not contain analogous data. However, this shortcoming affects only one of our quality measures (staffing levels).

We rely on a fourth source, Medicare's Dialysis Facility Compare website (and archives dating back to 2007), for facility addresses. We require these addresses in order to place facilities into geographic markets. Although addresses are included in the Cost Reports (but not the DFRs), they sometimes pertain to corporate headquarters. [35] Appendix 1 contains details pertaining to data cleaning.[36]

Our final sample pools both hospital-based and freestanding facilities. While prior research has focused exclusively on the latter, perhaps due to the challenges associated with obtaining data for hospital-based clinics, our interviews with industry experts suggested that both clinic types compete in the same product market. However, for robustness we also estimate all models excluding hospital-based facilities.

Figure 1: Four-Firm Concentration Ratios, U.S. Dialysis Industry 1996-2009

Figure 2: Sample page from Dialysis Facility Reports
Figure 3: Change in Local Market HHI, 2000-2009
Figure 4: Merging Clinic Locations
Figure 5: Predicted Impact of 2005/2006 Mergers ("Simulated Change in HHI")

[34] In 2011, ProPublica launched a user-friendly online tool which provides several key quality measures from the DFRs, including rates of mortality, hospitalization, and infections, as well as clinical benchmarks and a summary of inspection results.

[35] For clinics that do not appear in these data (primarily owing to closure prior to 2007), we utilize addresses reported in the Dialysis Cost Reports (for freestanding facilities) and the Hospital Cost Reports (for hospital-based facilities). Dialysis Facility Compare is a more reliable source of addresses than the Cost Reports, as the latter often use the corporate headquarters address. For hospital-based clinics, we use the hospital's main address.

[36] Roughly 10 percent of facility-years present in the Dialysis Cost Reports and Hospital Cost Reports lack accompanying DFR data, however most appear to be entering or exiting the market. For example, the USRDS reports 5,760 clinics in 2009. There are 5,703 observations in the DFRs. After dropping facilities which are recorded as not yet opened, already closed or having no data for any fields, there are 5,448 observations. After merging with other data sources and keeping only those facilities for which we can obtain an address, there are 5,410.

Figure 6: Relationship between Local Market Concentration and Dialysis Quality (Reduced Form Estimates)
Figure 7: Effect of Mergers on Chain Quality
Figure 8: Dialysis Revenue per Treatment (Quarterly Reports)
Figure 9: Private Dialysis Prices (HCCI)
Figure 10: Reduced Form Relationships between ΔHHI and Prices

Table 1 reports detailed summary statistics for our final dataset. Column 1 includes data from all years (i.e. 2000-2009). Columns 2 and 3 report data separately for 2005, the year immediately preceding the two major acquisitions we utilize for identification, and 2009, the last year in our sample. Figure 1: Four-Firm Concentration Ratios, U.S. Dialysis Industry 1996-2009

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Figure 7: Effect of Mergers on Chain Quality
Figure 8: Dialysis Revenue per Treatment (Quarterly Reports)
Figure 9: Private Dialysis Prices (HCCI)
Figure 10: Reduced Form Relationships between ΔHHI and Prices

Table 1 includes data from all facilities, excluding those with >50 percent pediatric patient population.

We estimate all models using five dependent variables. Three are measures of patient outcomes: the standardized mortality rate (SMR); the unadjusted death rate (expressed as the number of deaths per hundred patients); and the share of patients with URR greater than 0.75. Two are measures of inputs: the number of clinical staff members per patient,[37] and the share of patients with high hemoglobin levels. Key control variables include the percent of patients on Medicare, percent female, percent black, and average number of years with ESRD.

The SMR is defined as the actual number of deaths divided by the expected number of deaths; the expected number of deaths is a predicted value obtained from a model estimated by the Kidney Epidemiology and Cost Center using detailed individual-level clinic data. Until the 2007 DFR, the model did not include a control for institutionalized patients. Unfortunately, nursing home patients are non-randomly distributed across markets and firms, hence we limit our

[37] This is defined as the number of fulltime equivalent technicians, nurses, doctors, nutritionists and social workers per patient; administrators, janitorial staff, and other miscellaneous non-clinical workers are excluded. Similar measures are commonly utilized in the health services research literature. (Studies vary in which categories of workers are included; there is no general consensus and our results are insensitive to alternative definitions.)

analysis of SMRs to the time period during which we have a consistent time-series of data, 2003-2009.

While the SMR is recalibrated annually so that it averages roughly 1.0 in each year, the unadjusted death rate declined substantially over time, from 23.8 per hundred patients in 2000 to 20.9 per hundred patients in 2009.[38] The average number of years since ESRD diagnosis increased accordingly, from 3.7 to 4.3. URR ratios have also climbed.

Although hemoglobin levels depend on the interaction of patient characteristics, patient compliance, physician's orders, and facility influence, excessively high hemoglobin is a marker for overutilization of ESAs (which were lucrative to administer throughout our study period) – and hence we view this as a measure of inputs. Interviews with medical directors suggest that dialysis facilities have influence over the amount of ESAs which are prescribed, notwithstanding the fact that these drugs are sometimes ordered by physicians without direct financial links to the facilities. This anecdotal evidence is consistent with Hirth et al. (2009, 2010), who find that more of the variation in ESA levels is explained by facility fixed effects than physician fixed effects.

B. Dialysis Price Data

We obtain data on private prices for dialysis using the Health Care Cost Institute database (HCCI) from 2002-2011. HCCI's data contributors are Aetna, Humana and UnitedHealthcare,

[38] The SMRs reported in Figure 1: Four-Firm Concentration Ratios, U.S. Dialysis Industry 1996-2009
Figure 2: Sample page from Dialysis Facility Reports
Figure 3: Change in Local Market HHI, 2000-2009
Figure 4: Merging Clinic Locations
Figure 5: Predicted Impact of 2005/2006 Mergers ("Simulated Change in HHI")
Figure 6: Relationship between Local Market Concentration and Dialysis Quality (Reduced Form Estimates)
Figure 7: Effect of Mergers on Chain Quality
Figure 8: Dialysis Revenue per Treatment (Quarterly Reports)
Figure 9: Private Dialysis Prices (HCCI)
Figure 10: Reduced Form Relationships between ΔHHI and Prices
Table 1 do not average to one within each year for four reasons. First, prior to 2010, the normalization was done pooling four years of data at a time. (Since the 2010 DFRs contain data for 2006-2009, this normalization will impact the data years 2002-2005.) Second, the sample that KECC uses for risk adjustment is slightly different from the sample of facilities for which DFRs are produced. Third, we impose some sample restrictions. Fourth, we report un-weighted facility averages.

three insurers with national footprints. HCCI data contains the universe of claims for privately insured enrollees of these insurers. The claims include transaction prices, anonymized patient and facility IDs, and limited demographic information such as gender and age categories (0-17,18-24 and then 10 year increments until 85+), and zip code of residence.

We restrict our sample to patients who have 1-14 dialysis treatments in a month.[39] We further restrict the sample to adult patients with employer-sponsored insuranceand no prior transplant.[40] Nearly all ESRD patients without employer-sponsored insurance are Medicare eligible. Claims on behalf of patients with Medicare as a primary payer will be reimbursable at Medicare rates, which will be unaffected by the mergers. Pediatric and transplant patients have different spending levels and patterns which complicate price measurement.

Our ideal dependent variable is a price index for services provided by dialysis clinics to ESRD patients. Unfortunately, quantities are not consistently coded over time.[41] Furthermore, the set of commonly used procedure codes sometimes changes drastically from one time period to the next, leaving only a small share of revenue consistently coded over time.

We therefore rely upon two alternative measures of price, defined at the patient-quarter level: “narrow dialysis price”, which consists of quarterly payments for basic dialysis treatments divided by the quarterly number of these treatments, and “wide dialysis price,” which consists of quarterly outpatient facility and physician payments (i.e. no inpatient or prescription drug spending), also divided by the number of basic dialysis treatments.[42] Because quantities are not consistently coded over time, we infer the number of dialysis treatments per patient from the

[39] Only 0.13% of patients in our sample receives more than 14 dialysis treatments in a month. As we discuss in detail below, we infer the number of treatments someone receives from the dates on their claims. Therefore, someone with more than 14 treatments in a month usually has a claim with a nonsensical start or end date or dates spanning multiple months.

[40] HCCI ESI data contains a small sample of those with employer supplemental coverage for Medicare, akin to Medigap insurance. Because such a large share of the ESRD population is covered by Medicare, these individuals are a large share of the HCCI population on dialysis. We take four steps to eliminate these people. First, we throw out those with coverage types that are obviously supplemental. Second, we throw out anyone 65+. Third, we throw out individuals after they have been on dialysis for the 33 month period for which someone can keep ESI. Fourth, we throw out individuals after they have a physician claim reimbursed at 20% of the Medicare allowable amount for any of a number of relatively common procedures for which the quantity data appears accurate.

[41] The reported quantities are problematic for both dialysis treatments and for epoetin, the two largest cost components for these patients. As we discuss later, we develop and alternative method for determining the number of dialysis treatments that a patient receives in a month. We are unable to measure epoetin quantities consistently over time.

[42] Dialysis treatments are identified as claims with a procedure code of 90935, 90937, 90945, 90947, 90999, 90989, G0257, 99512 or revenue codes of 821, 831, 841, 851.

dates of service indicated on the patient's claims.[43] Dialysis and injectable drugs account for roughly 75% of the payments reflected in the wide price measure. The time series pattern from these measures is broadly consistent with the pattern from a chained Laspeyres Price Index, except that the Laspeyres Price Index has some additional sharp movements driven by coding changes.

Each measure has distinct advantages and disadvantages, which we discuss in turn. The narrow price measure exhibits few jagged movements, owing to the fact that the quantity of treatments and associated reimbursement is relatively cleanly measured over time, although we cannot definitively rule out the possibility that some bundled ancillaries are included. Dialysis treatments account for about half of the spending for our sample of patients, so the narrow measure is informative about overall spending. Regardless, it isn't clear whether one can extrapolate from the narrow price measure to the overall prices charged by dialysis facilities. In particular, within the Medicare population, relative margins across different procedures vary widely over time. It is therefore important to study a more inclusive price measure, in case margins for dialysis and ancillary services do not covary perfectly. In addition, total outpatient spending captures a number of high revenue codes which are associated with miscellaneous services, and which may not consistently be linked to dialysis services over time and across patients.

While our data contains de-identified facility IDs, due to confidentiality restrictions we cannot match facilities in the HCCI data to facilities in the Dialysis Facility Reports or Cost Reports data. Hence, we aggregate the HCCI data to the level of the HSA-quarter using patients' zip code of residence. Thus both our price variables and HCCI-derived demographic controls vary at the level of the HSA-quarter. We merge this data with HSA characteristics from the Dialysis Facility Reports and Cost Reports data. (Specifically, 2005 market shares for each chain and our measure of the merger's effects (i.e., ΔHHI, defined below)), Finally, we merge with county-year unemployment data, assigning each HSA to the county accounting for the plurality of its population.

Table 2 contains descriptive statistics for the HCCI data.

[43] More explicitly, one claim could cover a day, a week or a month of dialysis. The claim's quantity field does not reliably distinguish between these, but one can use the first and last date on a claim to infer the length of time covered by the claim and then, if one assumes the patient received three treatments per week (which is the usual amount), one can infer a number of treatments.

IV. The Evolution of Market Structure in the Dialysis Industry, 2000-2011

Because dialysis must be performed frequently (typically three times per week), and most patients face mobility challenges, geographic markets for dialysis are fairly small. For example, the Federal Trade Commission has relied upon county-level markets in its investigations of major acquisitions.[44] We rely on the Hospital Service Areas defined by the Dartmouth Atlas. Although the number of HSAs is comparable to the number of counties, HSA boundaries are determined using Medicare data on hospital choices, and will therefore more closely approximate geographic markets for dialysis.

The bottom panel of Figure 1: Four-Firm Concentration Ratios, U.S. Dialysis Industry 1996-2009

Figure 2: Sample page from Dialysis Facility Reports
Figure 3: Change in Local Market HHI, 2000-2009
Figure 4: Merging Clinic Locations
Figure 5: Predicted Impact of 2005/2006 Mergers ("Simulated Change in HHI")
Figure 6: Relationship between Local Market Concentration and Dialysis Quality (Reduced Form Estimates)
Figure 7: Effect of Mergers on Chain Quality
Figure 8: Dialysis Revenue per Treatment (Quarterly Reports)
Figure 9: Private Dialysis Prices (HCCI)
Figure 10: Reduced Form Relationships between ΔHHI and Prices
Table 1 reveals significant entry into new markets: the number of HSAs with at least one clinic increased by 20 percent between 2000 and 2009. By 2009, nearly 2/3 of the 3,436 HSAs contained at least one dialysis clinic. There was also significant expansion in existing markets. Among those HSAs with at least 1 clinic in 2000, 33.4 percent gained one or more clinics by 2009.

As noted previously, today more than two-thirds of clinics are operated by industry giants DaVita and Fresenius. Consolidation has largely taken place in a systematic, “food chain” fashion: independents have been acquired by small chains, small chains by large chains, and

[44] In each case, the FTC states "The relevant geographic market for the provision of dialysis services is defined by the distance ESRD patients are willing and/or able to travel to receive dialysis treatments, and is thus local in nature. Because ESRD patients often suffer from multiple health problems and may require assistance traveling to and from the dialysis clinic, these patients are unwilling and/or unable to travel long distances to receive dialysis treatment. As a general rule, ESRD patients do not travel more than 30 miles or 30 minutes to receive dialysis treatment, although travel times and distances vary depending on geographic barriers, travel patterns, and whether an area is urban, suburban, or rural."

large chains by the two industry leaders, DaVita and Fresenius. In addition, the largest chains have built new facilities at a fast clip.

Table 3 gives the details of the major acquisitions between 2000 and 2011, including the identities of the target and acquiring firms, the dates of announcement and completion, and the numbers of clinics involved.[45] All of the acquisitions were approved by the Federal Trade Commission, which typically required divestitures of select clinics so as to minimize the impact on local market structure. **Figure 3** is a histogram of the cumulative change in HHI for HSAs with at least one clinic as of 2000, excluding HSAs served by a single firm in both 2000 and 2009 (i.e. monopoly markets with no change in HHI).[46] The figure displays a good deal of variation in HHI changes, variation which will identify the coefficient of interest in our OLS models. Surprisingly, the figure also shows that the median change in HHI is negative. Thus, much of the industry consolidation over the past decade is associated with a broadening of the big two firms' geographic footprints (both through acquisition and de novo entry), rather than a deepening within markets in which they were previously active.

In Section VB below, we discuss the acquisition activity in greater detail. We will rely on variation in market concentration induced by large acquisitions to identify the causal impact of market concentration on quality.

V. The Impact of Local Market Structure on Dialysis Quality

A. Are Changes in Quality Correlated with Changes in Market Structure?

We begin our investigation by estimating simple models of the relationship between clinic-level dialysis quality and market concentration, as measured by the HHI. We estimate equations of the following form:

(1) $$\text{quality}_{\text{fmct}} = \beta \text{HHI}_{\text{m,t}-1} + \gamma_{\text{f}} + \tau_{\text{t}} \left[+\text{X}_{\text{fcm,t}-1}\boldsymbol{\lambda}\right] [+\nu_{\text{ct}}] + \varepsilon_{\text{fcmt}}$$

[45] The table reports the number of clinics as given in press releases. These figures correspond very closely with the numbers identified in our sample. The two major acquisitions we consider – of Gambro and Renal Care Group – reportedly involved 565 and ~450 clinics, respectively; in our data the corresponding totals are 570 and 474.

[46] Of the 1,786 HSAs with facilities in both 2000 and 2009, 983 were monopoly markets in both years.

where *f* denotes facility, *c* denotes chain, *m* denotes market, and *t* denotes year. In our baseline model, we regress quality on lagged market-area HHI, facility fixed effects (γ_f), and year fixed effects (τ_t). The coefficient of interest (β) is therefore identified by changes in market-level HHI.

Next, we add (lagged) covariates from the DFRs (X_{fcmt}): percent in Medicare, percent on hemodialysis, percent female, and average years on dialysis. These controls should explain some variation in quality measures (with the exception of the SMR, which already incorporates more detailed patient covariates), and could affect the coefficient of interest to the extent that changes in these covariates are correlated with changes in HHI. For example, if acquisitive chains tend to avoid Medicare patients, and Medicare patients are less healthy, we might expect the coefficient on HHI to be downward-biased when using the "observed death rate" as the measure of quality. Last, we include chain*year interactions (ν_{ct}), for the largest chains (>70 clinics during any year in the sample).[47] These terms allow for chain-specific variation in quality over time, and will be identified both by changes in quality following changes in chain affiliation, and by changes over time in the quality of specific chains.

We estimate all specifications by weighted least squares, using the facility end-of-year patient counts as weights. Standard errors are clustered by HSA. Our estimation sample excludes facilities in the 19 HSAs with more than 10 hospitals, as these HSAs are very broad (e.g., Houston, Chicago) and unlikely to be appropriate geographic markets for dialysis. We also drop facilities in which in-center hemodialysis is not the dominant treatment modality, and where pediatric patients account for more than half of the patient load.[48] Combined, these restrictions reduce the sample size by 17.5 percent.[49]

The OLS results are presented in **Table 4**. The estimates reveal no statistically significant association between concentration and quality, using any measure. Moreover, the

[47] Using 70 clinics as the cutoff (rather than, say, 100) ensures there are interactions for both of the chains which acquire clinics divested by Fresenius and DaVita in 2005/2006 pursuant Consent Orders issued by the Federal Trade Commission.
[48] When calculating HHIs, we drop pediatric facilities, but keep facilities in which in-center hemodialysis is not the not dominant treatment modality.
[49] Eliminating the largest cities reduces the number of facilities in our sample by 12 percent and the number of patients by 14 percent. The remaining restrictions eliminate an additional 5.5 percent of facilities, and 2.5 percent of patients.

magnitudes of the point estimates are very small: even the upper or lower bounds of the 95 percent confidence intervals imply small movements in quality in response to changes in market concentration. For example, consider a 500-point increase in HHI, approximately the change associated with a merger of two of six evenly-sized firms (assuming that all firms retain their pre-merger market shares after the merger). The estimated impact on the URR quality measure is a decrease of 0.06, as compared to a mean of 44. Using the lower-bound of the confidence interval implies a reduction of 0.19, also a trivial amount. The range of effects is modestly larger for SMRs (at the lower bound of the 95-percent confidence interval, a reduction of -.03 as compared to a mean of .99), and smaller for all other outcomes.

The lack of an association between changes in market concentration and changes in quality does not, of course, imply there is no causal relationship between the two. HHI is an endogenous variable, influenced by changes in patient preferences, firm strategies, and insurer decisions. For example, if larger chains within a given area offer better quality (perhaps due to economies of scale), then patients may flock to those clinics over time, resulting in a spurious positive association between quality and HHI. For these reasons, we turn to specifications that focus on plausibly exogenous changes in market concentration induced by major national acquisitions.

B. Is there a Causal Relationship between Clinic Quality and Local Market Concentration?

In this section, we attempt to estimate the causal effect of market concentration on quality by exploiting variation in local market concentration induced by M&A activity. By isolating variation due to structural changes, we eliminate the endogeneity bias on the HHI coefficient arising from factors such as changes in chain strategy or in the preferences of patients or their referring providers. However, local or regional M&A itself may be influenced by expectations of clinic quality trajectories. Hence, our empirical strategy is to focus solely on the *differential effect* of national acquisitions on different markets, owing to differences in pre-merger market shares of the target and acquiring chains. This strategy requires us to limit attention to sizeable mergers, i.e.

those listed in Table 3. Of the five mergers which close before the end of the study period, only the two largest involve non-trivial overlap in the locations of the target and acquiring chains.[50] The first of these is DaVita's acquisition of Gambro, which closed in October 2005. We estimate DaVita's facility count stood at 737 prior to the acquisition, and Gambro's at 570. **2009**

Figure 4 illustrates the locations of both firms. To address concerns about the most severe market overlaps (in particular those areas where the merger would create a monopoly), the FTC required divestitures of ~70 clinics; these were acquired by the newly-formed chain, Renal Advantage Inc. The second major transaction was Fresenius' acquisition of Renal Care Group, which closed in March 2006. By our estimates, the pre-merger facility counts were 1,135 and 474, respectively. The FTC required the newly-merged firm to divest ~100 sites, which were acquired by the newly-formed chain DSI. Interestingly, in 2011 DaVita acquired DSI, while in 2012 Fresenius acquired DaVita's spinoff, Renal Advantage.

We construct our merger-based instrument for HHI as follows:

(2)
$$\begin{aligned}\text{sim}\Delta\text{HHI}_{\text{m}} = &\sum(\text{market share of target}_{\text{m,2005}} + \text{market share of acquirer}_{\text{m,2005}})^2 \\ &- (\text{market share of target}_{\text{m,2005}})^2 - (\text{market share of acquirer}_{\text{m,2005}})^2 \\ &= \sum 2(\text{market share of target}_{\text{m,2005}})(\text{ market share of acquirer}_{\text{m,2005}}).\end{aligned}$$

$sim\,\Delta HHI_m$ represents the merger-induced *change* in market *m*'s HHI that would have occurred between 2005 and 2006 absent any other changes in market shares. **Figure 5** provides detail on the actual distribution of $sim\,\Delta HHI_m$ in the 91 HSAs in which it is non-zero. While $sim\,\Delta HHI_m$ combines the effect of both mergers, it is worth noting that there is little overlap in the markets affected by each transaction: there are 64 markets with non-zero $sim\,\Delta HHI_m$ due to the DaVita acquisition, 37 markets with non-zero $sim\,\Delta HHI_m$ due to the Fresenius acquisition, and 10 affected by both acquisitions.

[50] To be more precise, the number of HSAs in which both target and acquirer have locations is: Fresenius/Everest (9), RCG/NNA (5), DaVita/Gambro (84), Fresenius/RCG (76), Renal Advantage/National Renal Alliance (2). In addition, the Renal Advantage/National Renal Alliance merger closed at year-end 2008, implying a very short post-merger period of observation given our sample ends in 2009.

To mitigate the reduction in competition resulting from these acquisitions, the FTC required a substantial number of divestitures in each case. Both target and acquirer clinics were divested. Therefore, in some markets $sim\Delta HHI_m$ is negative.[51] We drop the 3 outlier markets with $sim\Delta HHI_m$<-2,500; these are cases in which the HSA is a poor approximation for actual geographic markets (i.e., there are clinics located near one another but in different HSAs). As in the OLS analysis, we also exclude HSAs with 10 or more hospitals.[52] The resulting sample includes 72 HSAs, with a mean $sim\Delta HHI_m$ of 1307. None of the results is sensitive to the sample restrictions.

B.1 First Stage Models

We propose to use $sim\,\Delta HHI_m * year > 2006_t$ as an instrument for $HHI_{m,t-1}$ in equation (1) above. To evaluate whether this instrument is indeed predictive of changes in HHI, we estimate the following model using market-year data, focusing exclusively on markets with non-zero $sim\,\Delta HHI_m$.

$$\text{(3)} \qquad \mathrm{HHI}_{\mathrm{mt}} = \lambda_{\mathrm{m}} + \tau_{\mathrm{t}} + sim\Delta HHI_m * \tau_{\mathrm{t}} + \varepsilon_{\mathrm{mt}}$$

The terms λ_m and τ_t represent a full set of market and year fixed effects, respectively. The results from this specification, presented in **Table 5**, reveal three key facts: (1) movement in HHI was largely uncorrelated with $sim\Delta HHI_m$ until 2006; (2) in 2006, the coefficient on $sim\Delta HHI_m$ increased by 0.88 ; (3) this increase attenuated only slightly during 2007-2009. (Note that one would expect an increase in the coefficient on $sim\Delta HHI_m$ between 2005 and 2006 to equal one only if three conditions are satisfied: (1) market shares for every clinic are completely unaffected by ownership changes; (2) there are no other sources of fluctuations in market shares over time; (3) there is no measurement error.) We conclude the major acquisitions of 2005/2006 did indeed

[51] Note we assume divestitures are also orthogonal to changes in the outcomes of interest, i.e. that divestiture locations are not systematically correlated with future shocks to quality. Given the FTC used clear guidelines pertaining to HHI and chain presence in order to identify markets requiring divestitures, this assumption seems relatively weak (and particularly so conditional on the interaction terms which will be included in the model). A list of divested clinics is located in the Consent Orders from each FTC case (see http://www.ftc.gov/opa/2005/10/DaVita.shtm and http://www.ftc.gov/opa/2006/03/fresenius.shtm for details).

[52] The mean value for $sim\,\Delta\,HHI_m$ in the HSAs with 10+ hospitals is 806.

generate large shocks to local market structure, shocks which were uncorrelated with prior trends in market-specific HHI and which persisted for a significant period following the merger.

While the results from equation (3) provide support for the natural experiment underlying our 2SLS model, Table 5 does not represent the first stage of this analysis. The first stage must be estimated using the facility-year data, and including additional controls; the results (presented in **Appendix Table1**) closely mirror the findings in Table 5.

B.2 Reduced Form Models

To examine the effect of the merger-induced shocks to market concentration on quality, we estimate the following equation, separately for each quality measure:

$$\text{(4)} \qquad \text{quality}_{\text{fmct}} = \text{sim}\Delta\text{HHI}_{\text{m}} * \tau_{\text{t}} + \gamma_{\text{f}} + \tau_{\text{t}} \; [+\text{X}_{\text{fcm,t}-1}\boldsymbol{\lambda}] \, [+\delta_{\text{target,2007+}}][+\nu_{\text{ct}}] + \; \varepsilon_{\text{fcmt}} \, .$$

The controls in this model are identical to those utilized in the fixed-effects model (equation 1), however in equation (4) we consider a new interaction term, denoted $\delta_{target,2007+}$. This term captures the direct effect of acquisition on facility quality.[53] Once included, the coefficients of interest should reflect the effects of simΔHHI on quality, controlling for the effect of ownership changes on quality. (For example, suppose that DaVita boosted the quality of Gambro clinics after acquisition, but this increase was attenuated in areas where DaVita gained a greater market advantage. Controlling for the absolute change in quality will help to separate out these effects.

An alternative approach is to eliminate facilities owned by the target and acquiring firms; we discuss these results below.) Last, we add chain*year interactions (as before, for chains with >70 clinics). These interactions help to mitigate additional concerns about omitted variables bias (e.g., they allow Fresenius to have different annual changes in quality, which might otherwise be captured in the coefficient on simΔHHI because greater Fresenius presence in a market is correlated with simΔHHI). While these interactions might absorb some non-spurious variation in the outcome measures, both acquirer/target pairs have a substantial number of clinics which

[53] Note this control assumes a one-year lag in quality changes following an acquisition; more flexible controls do not affect the results, and we present findings with and without this control.

do not experience any change in competition from the merger and therefore aid in separately identifying these coefficients.

We estimate equation (4) by weighted least squares, using end-of-year patient counts as weights, and clustering the standard errors by HSA. **Figure 6** displays the estimates of the coefficients on $\text{sim}\Delta\text{HHI}_{\text{m}} * \tau_{\text{t}}$, along with 95 percent confidence intervals. In the interest of parsimony, the graph includes only the estimates from the full model with all bracketed terms. Recall that 2006 is a transition year (the first year following the DaVita acquisition, and the year of the Fresenius acquisition), so that the "post" treatment period truly begins in 2007. The graphs reveal no pre-merger trends in three of the five measures: SMR, observed death rate, and clinical staff per patient. However, clinics in markets more heavily affected by the merger appear to experience increases in % hemoglobin>12 and %URR>0.75. This pattern is consistent with heavy epoetin use and URR-boosting processes known to be implemented by large chains.[54] We therefore add linear market trends to our reduced-form specifications, although the results are largely insensitive to their inclusion.

With one exception, there are only modest changes in post-acquisition coefficients. The exception is the share of patients with high hemoglobin levels. Table 6 presents regression coefficients from a parsimonious version of equation (3), i.e. replacing $\mathbf{sim\Delta HHI_m} * \boldsymbol{\tau}_{\mathbf{t}}$ with $\mathbf{sim\Delta HHI_m} * (\mathbf{year} > \mathbf{2006})_{\boldsymbol{t}}$. This table reveals no statistically significant deviations from trend during the post-merger period. All point estimates are very small, and the signs do not consistently point to higher or lower quality. Even the hemoglobin surge apparent in   is rather small (and, as previously stated, statistically insignificant). The point estimate for the coefficient on $\mathbf{sim\Delta HHI_m} * (\mathbf{year} > \mathbf{2006})_{\boldsymbol{t}}$ is approximately 10, implying a 1000-point increase in simΔHHI would increase the percent of patients with excess hemoglobin by 1 (compared to a mean of 37).

To gain a deeper understanding of the competitive response to DaVita and Fresenius acquisitions, we estimate the following model

(5) $\text{quality}_{\text{fmct}} = \beta_1 \text{sim}\Delta\text{HHI}_{\text{m}} * \text{merging}_{\text{mc}} + \beta_2 \text{sim}\Delta\text{HHI}_{\text{m}} * \text{non} - \text{merging}_{\text{mc}} + \gamma_{\text{f}} + \tau_{\text{t}}$

[54] Large chains employ a variety of techniques to boost clinical outcomes, including programs to encourage usage of fistulas, which permit more effective dialysis. Technically, the chain*year interaction terms should absorb the effects of these practices for the large chains, however there appear to be spillover effects on smaller rivals.

$$+ \rho_{mt} \, [+X_{fcm,t-1}\boldsymbol{\lambda}] \, [+\delta_{target,2006+}][+\nu_{ct}] + \varepsilon_{fcmt},$$

where "merging" is a market and chain-specific indicator which takes a value of 1 if the chain is either a target or an acquirer in market m, "non-merging" is the complement of this indicator, and ρ_{mt} are the linear market trends previously discussed. The results, presented in **Table 7** showed no consistent or meaningful differences across merging and non-merging facilities.

Our findings are robust to a variety of modifications to our model and estimation sample, including (1) using fixed-radius circular markets in place of HSAs; (2) excluding hospital-based facilities; (3) including facilities in markets with no merger -induced changes in HHI; (4) adding back in facilities in the largest HSAs; and (5) evaluating each of the mergers separately. In light of the reduced form results, we do not present instrumental variables estimates of the relationship of interest.

B.3. Extension: Is Quality a Local Choice Variable?

There are a number of possible explanations for the non-response of local quality to changes in local market structure. One possibility is that patients and/or their referring providers are not highly responsive to quality. If they were, and if local quality adjustments were not too costly, then for-profit chains would respond to the market incentive created by weakened competition. Such unresponsiveness to quality may be due to insufficient publication of quality data, and/or to other factors which impede the movement of patients and physicians across clinics.[55] Patients select clinics based on convenience and the advice of their nephrologists. Once they have established a relationship with a dialysis provider, switching facilities is unappealing due to significant transport challenges as well as personal relationships with dialysis technicians. Physicians prefer to round at a small number of clinics (ideally one), and if they serve as a

[55] In contrast, hospital choice for cardiac surgery is responsive to quality report cards (Dranove and Sfekas, 2008). There are a number of possible reasons for the disparity in quality responsiveness. First, dialysis patients are socioeconomically disadvantaged and tend to face cognitive challenges that preclude aggressive investigation of treatment options. Second, physical location is much more important for a recurring service. Third, the hospital industry has contracted over the last 30 years (at least in terms of the number of beds), while the dialysis industry has expanded dramatically. Exit from the hospital industry is likely to cull low quality providers, regardless of whether cardiac patients respond to quality information.

medical director for a clinic they often commit to a ten-year contract. Thus, shifting across clinics based on quality is difficult and slow.[56]

A second possibility is that chains may not separately adjust the quality of each facility in response to local competitive conditions. Chains may select a set of standardized procedures and simply work to implement them in all acquired facilities as quickly as is feasible and economical.[57] In many other industries, the ability of chains to deliver standardized quality across markets generates value for risk-averse consumers. This value proposition is divorced from local competitive conditions.

If quality is a national decision, we would expect to observe reductions in the quality of merging chains (and their close rivals) in the wake of the large mergers. We consider this possibility empirically by estimating regression models of the following form:

$$\mathrm{SMR}_{\mathrm{fmct}} = \gamma_{\mathrm{f}} + \tau_{\mathrm{t}} + \nu_{\mathrm{c2005}} * \tau_{\mathrm{t}} + \varepsilon_{\mathrm{fcmt}} .. \qquad (6)$$

Results, plotted in **Figure 7**, suggest that SMRs worsened at one of the acquired chains (RCG) and improved at the other (Gambro). The acquiring firms show no statistically significant change. We can rule out increases in SMRs above 0.02 for all merging chains from 2006 onward at the 95% confidence level.

If quality decisions are strategic complements, then including competitors of the merging firms in our control group will bias our estimates towards zero. We performed two additional analyses to confirm the robustness of the above results. First, if closer competitors of the four chains decreased quality, it stands to reason that markets with a higher pre-merger market share of the four chains would experience greater quality reductions post-merger. However, **Table 8** suggests that changes in SMRs do not depend upon the pre-merger market share of these four firms. Next, we compare the time trend for facilities owned by one of the four merging chains in 2005 to the time trend for non-chain facilities *in markets with no chain presence* in 2005 (i.e. facilities which experienced no change in competition from the mergers). The results (available upon request) closely mirror those using all facilities as a control group, but are less precisely

[56] These insights were culled from interviews with medical directors across a range of dialysis chains, as well as meetings with a regional director and a facility administrator for one of the market leaders.

[57] The management of one clinic we visited explained that the facility would eventually transition to the owner's preferred policy regarding dialysis filters, but was retaining the old system at present due to the cost of the switchover.

estimated. We can rule out increases in SMRs above 0.04 for all merging chains from 2006 onward at the 95% confidence level.

Together these results suggest that merger-induced reductions in competition were not followed by meaningful reductions in national quality. We are unable to rule out the possibility that simultaneous quality improvements (e.g., from increased economies of scale) offset quality reductions from decreased competition. However, from a policy perspective, our results provide no evidence that these mergers worsened the quality of care.

V. The Impact of Local Market Structure on Dialysis Prices for Privately-Insured Patients

We begin by presenting general trends in dialysis revenues, as reported by public sources. **Figure 9** plots quarterly dialysis revenue per treatment, as reported in quarterly 10Q filings by DaVita, Gambro, Fresenius and RCG. (Publicly traded companies are required by the SEC to file a 10Q every quarter. 10Qs are similar to a company's annual report, but typically less detailed and unaudited.) DaVita's revenues per treatment increased following the acquisition of Gambro in October 2005. Fresenius' surge in revenues per treatment predated its major acquisition (of RCG), but the trend appears to accelerate in the years immediately following. From 2007 to 2011, Davita's price per treatment remains relatively stable at the new, higher level, whereas Fresenius's surges again from mid-2008 until 2010. Of course, these aggregate increases cannot be separated from secular time trends driven by omitted factors, such as the cost of inputs or changes in services delivered. In particular, injectable spending within the Medicare population changes substantially over this time period, increasing in the earlier years, before decreasing starting around 2005-2006. Furthermore, Medicare introduced a new reimbursement scheme in 2011 which may have had spillover effects on privately-insured patients.

Before turning to those models, it is worth comparing the aggregate trends from the 10Q reports to the narrow and broad price measures constructed using the HCCI data. For purposes of confidentiality, **Figure 9** presents HCCI price measures normalized by the narrow price of dialysis in 2002Q1. There are small increases in both the narrow and wide price measures directly after/concurrent with the mergers. Both price measures are then relatively flat for about a year (2006Q4 to 2007Q3), at which point there is a sudden drop in revenues per treatment.

That drop is much more pronounced in the wide price measure, but is present in both. Further investigation suggests that it is largely attributable to decreased spending on injectables. Our understanding is that a bundled payment scheme was implemented by one or more private payers at that time, and that much of the drop is caused by a decrease in utilization rather than price. The drop also appears in the narrow price measure, albeit to a much smaller extent. This suggests either that some of the drop is in fact a price decrease or that some ancillaries are included in the narrow price measure, notwithstanding our efforts to exclude them. As previously noted, a similar drop appears in a price index. The takeaway is that we will be cautious in interpreting our estimates, as our price measure (particularly the wide version) appears to reflect utilization to at least some degree.

Immediately following the price drop at the end of 2007, dialysis prices begin a steep, steady increase that persists until the end of the study period. This trend begins a bit later than one might expect for a merger-induced change, however given infrequent contract negotiations this trend may be merger-related.

Turning to our regression analysis, we note a few differences from the quality analysis. First, the price analysis uses a slightly different level of aggregation of the key independent variable (ie. ΔHHI, the merger-induced *change* in market *m*'s HHI that would have occurred between 2005 and 2006 absent any other changes in market shares) than the quality regressions. For the quality regressions, ΔHHI is calculated at the HSA level. For the pricing regressions, we aggregate ΔHHI to the LEHID (Large Employer Health Insurance Dataset) market level, using a private treatment weighted average of the ΔHHIs for the HSAs within that LEHID market. The reason for the alteration is that provider prices are negotiated at more aggregated levels of geography than the HSA. The 139 LEHID markets are delineated by the insurance industry and reflect the geographic boundaries insurers use when quoting premiums.[58] For details see Dafny (2010). Specification (7) below uses ΔHHI_{m_l}, where the *l* subscript denotes the LEHID market definition.

A second difference from the quality specification is that we do not have facility identifiers and therefore cannot control for facility ownership. Instead, we create 4 variables: *% Davita 2005, % Gambro 2005, % Fresenius 2005* and *% RCG 2005*, which contain HSA-level

[58] Most are major metropolitan and ex-metropolitan areas in the same state, e.g. Chicago-area, Northern Illinois – excluding Chicago, Southern Illinois. For details see Dafny (2010).

shares of each of the four merging chains as of 2005. We include these variables in some specifications, so as to distinguish between chain-level price changes and price changes that are correlated with local changes in the level of competition. For consistency with the quality results, we refer to these shares as $\nu_{m_h ct}$ for HSA (m_h), chain (c) and quarter (q). We subscript m by *h* to emphasize that these variables are calculated at the more granular level of the HSA. We therefore run the regressions:

$$\text{(7)} \qquad \ln(\text{Price}_{m_h t}) = \Delta\text{HHI}_{m_l t} * \tau_t + \gamma_{m_h} + \tau_t \ [+X_{m_h,t}\boldsymbol{\lambda}]\ [+\nu_{m_h ct}] + \ \varepsilon_{m_h t}$$

where $X_{m_h,t}$ are market-time controls (specifically lagged annual unemployment and HSA-quarter demographics from the HCCI sample: shares in different age groups (<45, 45-54, 55+); % male; and % HMO).

The results are displayed in **Figure 10**. We plot τ_t, the interactions between quarter dummies and $\Delta\text{HHI}_{m_l t}$ for each of the two mergers separately. Positive coefficients after the merger imply price increases in areas with merger-induced increases in local market concentration. The coefficients in Figure 10 derive from specifications that do not control for the 2005 chain shares, however. Thus, these coefficients also capture chain-level changes in price to the degree they are correlated with our instrument. The results are broadly similar when those interactions are included (and will appear in a later version of the paper after another round of HCCI disclosure reviews).

For our narrow price measure (presented in Figure 10, Panel A), the results are broadly consistent across the two mergers. There is a noisily-estimated negative coefficient on $\Delta\text{HHI}_{m_l t}$ until 2009, when the coefficient turns positive and begins a steady increase. Given the mergers occur in late 2005/early 2006, the increase might be attributable, at least in part, to non-merger factors that are correlated with $\Delta\text{HHI}_{m_l t}$. We evaluated two potential non-merger explanations for the break in trend. First, the coefficients on $\Delta\text{HHI}_{m_l t}$ begin increasing around the same time as the large drop in the wide price measure. We consider the possibility that the increased rate of growth reflects compensation to dialysis providers for lower rents from drugs. A closer examination of the data suggests that the price increases from 2008 onwards are not localized to the markets affected by the move to bundled payments, however. A second explanation is that the change in pricing is related to the recession; the timing is also coincident with the recession's start. However, the results are insensitive to inclusion of the local unemployment rate. We are not aware of other candidate non-merger factors. However, the lagged response does temper our conclusions.

For our wide price measure (Figure 10, Panel B), the results are directionally similar, however there are greater differences across the two mergers. In particular, markets affected by the Davita/Gambro merger experience a sharp drop in price two years after the merger closed, coincident with the general drop illustrated in Figure 9 and (we believe) driven by the move towards bundling drugs and decreasing utilization. With the exception of that drop, the trend is also generally positive in the post period for that merger. The results for this merger effect are therefore particularly open to differing interpretations, as it appears the price series captures utilization changes. We hope to explore this issue further in future versions of this paper.

Table 9 presents results from a more a parsimonious specification that replaces the quarterly interactions with $\Delta\mathrm{HHI}_{\mathrm{m}_{\ell}\mathrm{t}}$ with interactions between $\Delta\mathrm{HHI}_{\mathrm{m}_{\ell}\mathrm{t}}$ and a post-period indicator. As with the quality results, we define the post period as beginning in 2007 (for both mergers). We also consider a specification that adds an interaction between post*$\Delta\mathrm{HHI}_{\mathrm{m}_{\ell}\mathrm{t}}$ and a linear time trend. The coefficient on this interaction will capture the upward trend in post-merger coefficients on $\Delta\mathrm{HHI}_{\mathrm{m}_{\ell}\mathrm{t}}$ depicted in Figure 10. All columns include four controls capturing the 2005 share of each merging chain interacted with a post-period indicator. The specifications with the linear trend interactions also include interactions between these controls and a linear trend.

The results reveal no post-period impact of either merger on either measure of dialysis price, *on average*. However, the effect of the merger on both price measures increases steadily and significantly over time. Panel B presents implied coefficients (and standard errors) for the final quarter in the study period, 2011Q4. The estimate of 2.84 (p=.054) for the narrow price measure), multiplied by the mean of $\Delta\mathrm{HHI}_{\mathrm{m}_{\ell}\mathrm{t}}$ (129 points), yields 0.037. Thus, we predict that narrow dialysis prices were 3.7% higher than they would have been in the absence of the mergers. The analogous figure for the wide price measure is 2.5% (p=.094).

VI. Conclusions

Over the past decade, the U.S. dialysis industry has nearly doubled in size. At the same time, the industry has consolidated into a for-profit duopoly, with multimillion-dollar executive pay - in spite of its largely Medicare-insured population. These developments have prompted high-profile articles and investigations alleging a variety of misconduct, ranging from the overuse of the highly-reimbursed anti-anemia drug Epogen (and its biosimilars) to the flouting of safety regulations. Most importantly, patient survival rates and quality of life is considerably lower in the U.S. than in other developed countries.

In this study, we explore the extent to which the industry consolidation has diminished quality competition. We find no evidence that consolidation within local markets affects the quality of care in those markets, or the use of inputs (specifically staff and epo). Furthermore, the national quality of the merging chains relative to a number of different control groups did not decrease after the mergers.

One possible explanation for our results is that consumer demand is unresponsive to quality in this industry, which limits the incentive for facilities to improve quality. If true, payers and public agencies may wish to consider policies to strengthen the link between competition and quality, for example by aggressively disseminating quality information to consumers, and/or by re-orienting published quality information to shame low-quality firms (Ramanarayanan and Snyder 2012). "Pay-for-performance" is another alternative. Indeed, CMS has moved in this direction, implementing a "Quality Incentive Program" which reduces reimbursements for centers with low measured quality.[59] More aggressive facility inspections (with teeth) are yet another possibility.

A second possible explanation for the lack of adverse quality effects in markets exposed to post-acquisition increases in concentration is that the Federal Trade Commission's local divestiture requirements prior to each acquisition were effective (at least on this dimension). That is, either our measures of market structure are inaccurate, and/or any increases remaining after

[59] By 2014, the Quality Incentive Program will include thresholds for dialysis adequacy, hemoglobin levels, and fistula access. To earn full reimbursement, facilities must also report dialysis-related infections, administer patient satisfaction surveys, and monitor phosphorus and calcium levels on a monthly basis. (Source: *Renal Business Today*, 11/2/2011, available at http://www.renalbusiness.com/news/2011/11/cms-raises-dialysis-reimbursement-for-2012-revises-qip.aspx.)

the divestitures are not material. Given we find some evidence of price effects, we lean toward quality inelastic demand as the likelier explanation for the lack of a relationship between competition and quality.

We also explore whether market consolidation led to higher prices. While Medicare reimbursements are fixed, private insurers negotiate prices with dialysis facilities. Thus far, we find some evidence that merger-induced increases in local market concentration are correlated with higher private prices. These market-specific increases in price, however, pale in comparison to national increases in price. It is difficult to link these increases conclusively to the large industry mergers we study, both due to the absence of a control group and due to a shift toward bundled payment several quarters following the merger.

In closing, we find that local quality and price competition in the dialysis industry appears to be limited. Competition, particularly on the pricing side, may take place nationally or regionally, notwithstanding the fact that services are consumed locally. Additional insights could be gained by examining whether this finding generalizes to sectors with similar characteristics, such as nursing homes and rehabilitation centers.

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Appendix 1

This appendix provides additional details about the construction of our dataset. Our primary source of data are the Dialysis Facility Reports. In constructing our dataset, we had to deal with three complications, which we now cover in turn. First, we had to decide which years of information to retain from each DFR. Second, we had to make some adjustments to track the same facility over time. Third, we merged the DRFs, Cost Reports and Dialysis Facility Compare databases. This is complicated by the presence of up to three provider numbers on each DFR.

We retained only the most recent vintage of each Dialysis Facility Report. On might worry that this will lead to missing final years of data for facilities that exit. This is not the case - the DFRs are produced for four years after a facility closes. In cases in which we would obtain extra years of data by choosing an alternative vintage for a facility (ie in cases in which a facility appears to have a missing DFR), we were able to ascertain that in a vast majority of cases, the cause was a change in the facility's Medicare provider numbers, and therefore that the mixing vintages of data would result in double counting some facilities.

If one DFR and one observation in the Cost Reports were always associated with just one facility, then tracking facilities over time would trivial. This is not the case, so we discuss the steps necessary to track facilities. Each DFR can list up to three Medicare provider numbers, the dialysis operations of which are then consolidated onto one Cost Report. The most common reason for this is that the DFR contains multiple provider numbers associated with one hospital. Sometimes, multiple provider numbers exist because of a change in a facility's provider number (ie, we observe the simultaneous entry of one provider number and the exit of another). We use these features to track the same provider over time. In a small number of cases in which the set of provider numbers on a DFR has changes inconsistent with tracking the same facility over time, we allow for breaks in tracking the facility, as these cases represent occurrences like mergers. These complications affect a very small number of (predominantly hospital based) facilities, and our results are robust to the exclusion of hospital based facilities.

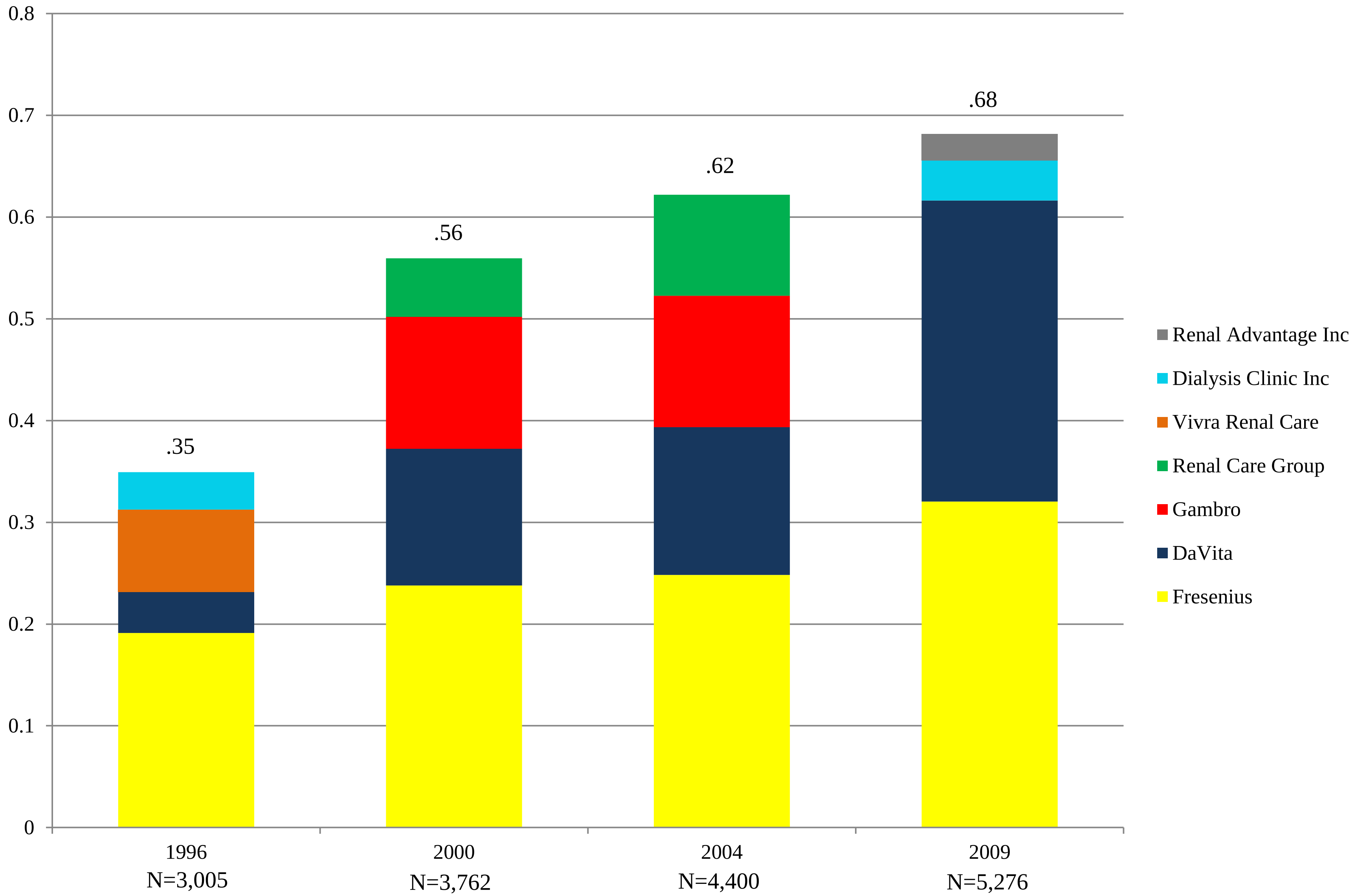
Next, we merged the Cost Report data and the Dialysis Facility Compare data into our dataset of DFRs. Recall that one DFR can have multiple provider numbers. Furthermore, one hospital can have more than one outpatient dialysis facility. Therefore, there is potential for non-unique matching in both directions. To choose a unique match for each DFR, we give priority to hospital satellites over hospitals (ie if one DFR has the provider number of a hospital and a hospital satellite dialysis facility, we use the hospital satellite dialysis facility's address)[60], and to the first provider number listed on a DFR over the second or third. These restrictions uniquely match each DFR to at most on one observation in the Cost Reports and one observation in Dialysis Facility Compare. Again, these complications primarily affect hospital facilities (ie. if one hospital has two outpatient facilities, the hospital's provider number may appear on two DFRs).

Our final dataset retains all DFRs for which we are able to obtain an address, which is over 99% of our DFR dataset.[61] We cannot judge how well our merging works based on how many hospitals are unmatched because most hospitals do not have dialysis facilities. For freestanding dialysis facilities, about 1% of Cost Reports do not have a match in the Dialysis Facility Reports. About 75% of these observations are in the year that a facility enters or exits. Overall, this suggests that our merging procedures worked well.

[60] Because we do not use data from hospital Cost Reports, this only affects the facility address and not any other data.
[61] Because we have address data from both the Cost Reports and the Dialysis Facility Compare database, we obtained geocodes for both and checked the largest cases in which the geocodes differed.

Tables and Figures

Figure 1: Four-Firm Concentration Ratios, U.S. Dialysis Industry, 1996-2009



Note: N refers to the number of clinics. All statistics are unweighted. The unit of observation is the facility-year. The sample is all observations used in calculating HHIs, and therefore excludes facilities with >50% pediatric patients in any year.

Figure 2: Sample Page from Dialysis Facility Report

2009 Dialysis Facility Report

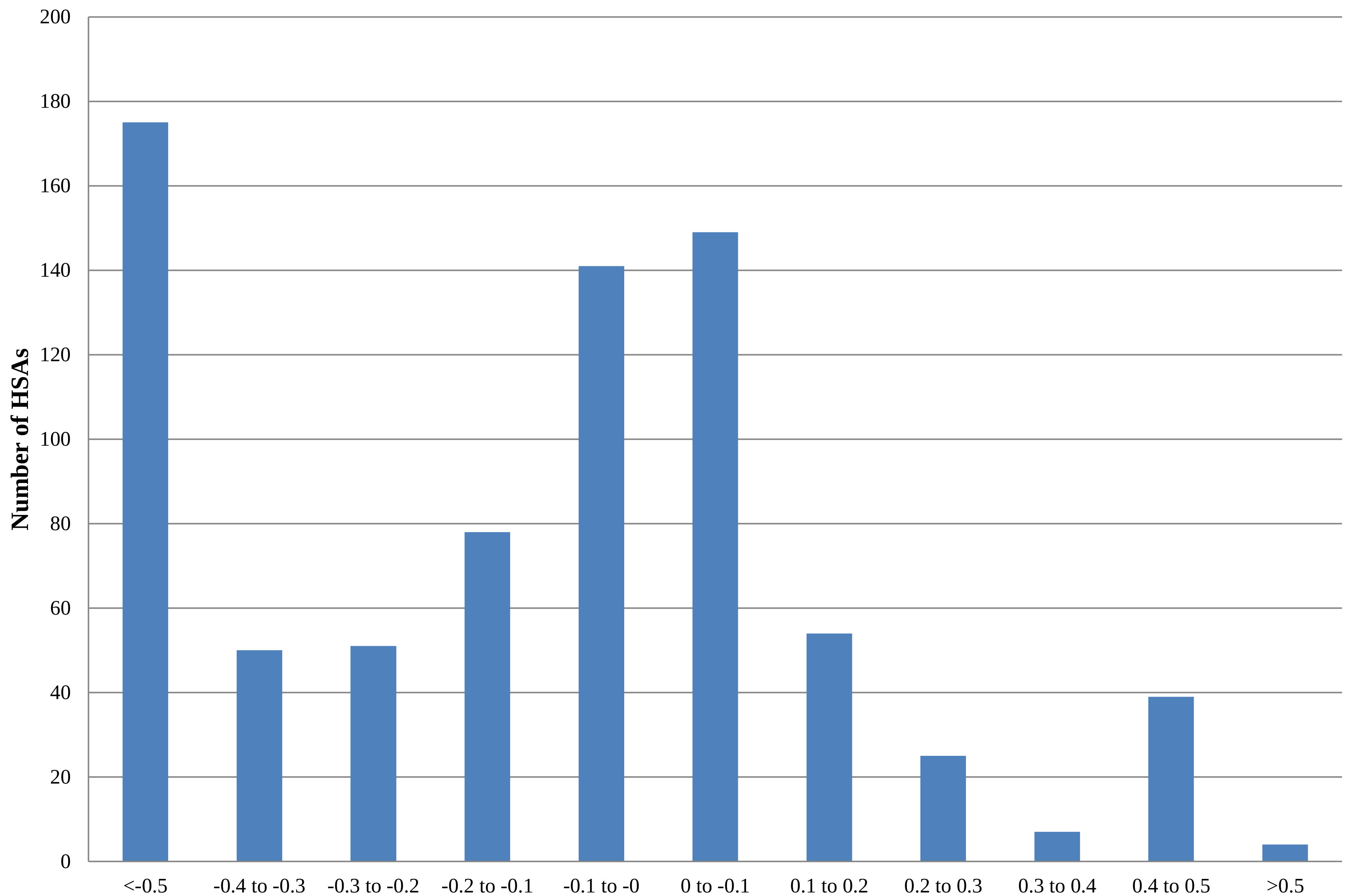
BIRMINGHAM EAST DIALYSIS State: AL Network: 08 CMS Provider#: 012508

TABLE 1: Mortality Summary for All Dialysis Patients (2005-08) & New Dialysis Patients (2005-07)[1]

		This Facility					Regional Averages[2], per Year, 2005-2008		
	Measure Name	**2005**	**2006**	**2007**	**2008**	**2005-2008**	**State**	**Network**	**U.S.**
All Patients: Death Rates									
1a	Patients (n=number)	140	155	143	141	579[7]	74.5	77.4	90.1
1b	Patient years (PY) at risk (n)	90.6	105.1	110.4	97.8	403.9[7]	53.5	55.3	60.6
1c	Deaths (n)	20	23	18	15	76[7]	11.4	11.6	12.7
1d	Expected deaths (n)	13.7	15.4	16.6	15.0	60.7[7]	10.2	10.7	12.7
1e	Death rate per 100 PY (% of 1b)	22.1	21.9	16.3	15.3	18.8	21.4	20.9	21.0
1f	Expected death rate per 100 PY (% of 1b)	15.1	14.7	15.1	15.3	15.0	19.1	19.4	21.0
All Patients: Categories of Death									
1g	Withdrawal from dialysis prior to death (% of 1c)	15.0	4.3	0.0	0.0	5.3	18.6	19.5	23.8
1h	Death due to: Infections (% of 1c)	5.0	0.0	11.1	20.0	7.9	14.5	14.7	17.5
	Cardiac causes (% of 1c)	15.0	0.0	11.1	13.3	9.2	23.4	23.7	27.3
1i	Dialysis unrelated deaths[3] (n; excluded from SMR)	0	0	0	0	0[7]	0.1	0.1	0.1
All Patients: Standardized Mortality Ratio (SMR)									
1j	SMR[4]	1.46	1.49	1.08	1.00	1.25	1.12	1.08	1.00
1k	P-value[5]	0.13	0.08	0.80	0.99	0.07	n/a	n/a	n/a
1l	Confidence interval for SMR[6]								
	High (95% limit)	2.26	2.24	1.71	1.65	1.57	n/a	n/a	n/a
	Low (5% limit)	0.89	0.95	0.64	0.56	0.99	n/a	n/a	n/a
All Patients: SMR Percentiles for this Facility (i.e. percent of facilities with lower mortality rates)									
1m	In this State	75	76	50	46	70			
1n	In this Network	77	80	56	50	75			
1o	In the U.S.	83	86	64	59	82			
							Regional Averages Per Year, 2005-2007[2]		
New Patients: First Year Death Rates		**2005**	**2006**	**2007**		**2005-2007**			
1p	New Patients (n=number)	41	43	24		108[7]	17.2	18.2	20.2
1q	Patient years (PY) at risk (n)	34.5	38.5	20.0		93.1[7]	14.6	15.5	17.1
1r	Deaths (n)	9	6	5		20[7]	4.1	4.4	4.8
1s	Expected deaths (n)	7.8	6.2	4.4		18.5[7]	3.8	4.2	4.8
1t	Death rate per 100 PY	26.1	15.6	25.0		21.5	27.9	28.2	28.1
1u	Expected death rate per 100 PY	22.7	16.2	22.0		19.8	26.2	27.3	28.1
New Patients: Categories of Deaths									
1v	Withdrawal from dialysis prior to death (% of 1r)	11.1	16.7	0.0		10.0	26.8	28.8	32.2
1w	Death due to: Infections (% of 1r)	0.0	0.0	0.0		0.0	19.7	19.0	23.1
	Cardiac causes (% of 1r)	11.1	0.0	20.0		10.0	25.9	28.4	33.1
New Patients: First Year Standardized Mortality Ratio (SMR)									
1x	SMR[4]	1.15	0.96	1.14		1.08	1.06	1.03	1.00
1y	P-value[5]	0.77	0.99	0.90		0.78	n/a	n/a	n/a
1z	Confidence interval for SMR[6]								
	High (95% limit)	2.18	2.10	2.65		1.67	n/a	n/a	n/a
	Low (5% limit)	0.53	0.35	0.37		0.66	n/a	n/a	n/a

n/a = not applicable [1] See *Guide, Section IV.*
[2] Values are shown for the average facility, annualized.
[3] Defined as deaths due to street drugs and accidents unrelated to treatment.
[4] Calculated as a ratio of deaths (1c to 1d for all patients, 1r to 1s for new patients) to expected deaths; not shown if there are too few expected deaths.
[5] A p-value less than or equal to 0.05 indicates that the difference between the actual and expected mortality is probably real and is not due to random chance alone, while a p-value greater than 0.05 indicates that the difference could plausibly be due to random chance.
[6] The confidence interval range represents uncertainty in the value of the SMR due to random variation.
[7] Sum of 4 years (all patients), or 3 years (new patients), used for calculations; should not be compared to regional averages.

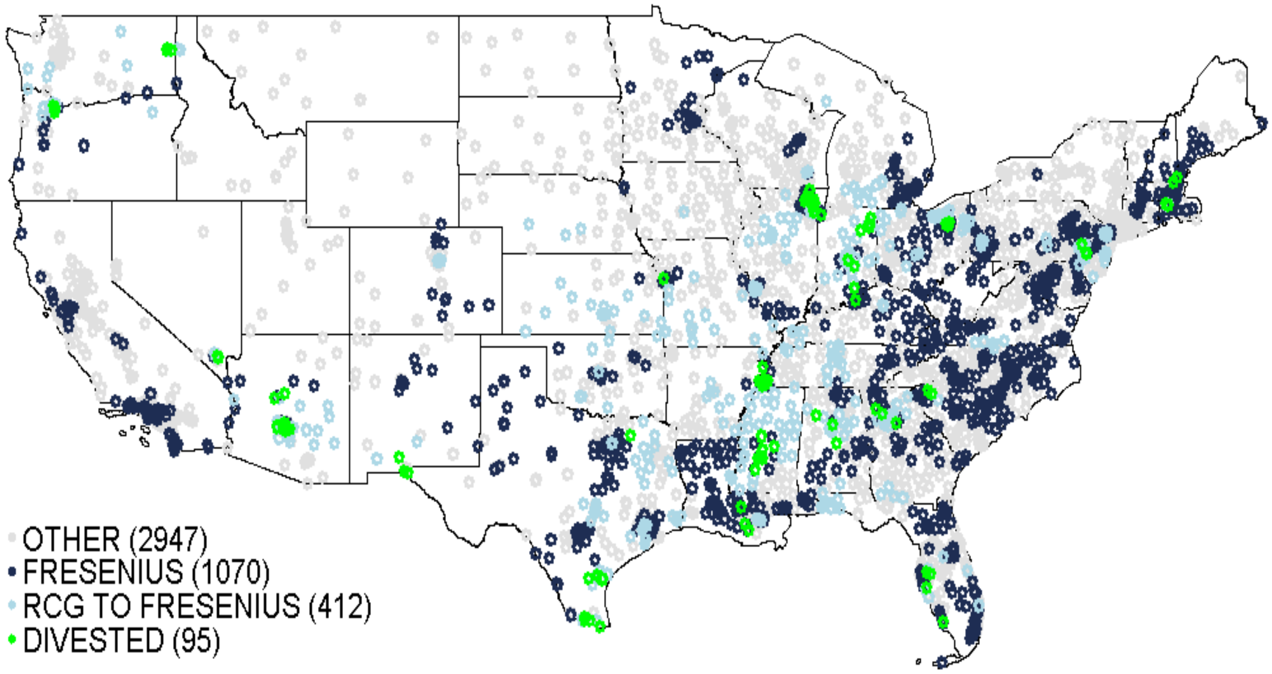
Figure 3: Change in Local Market HHI, 2000-2009



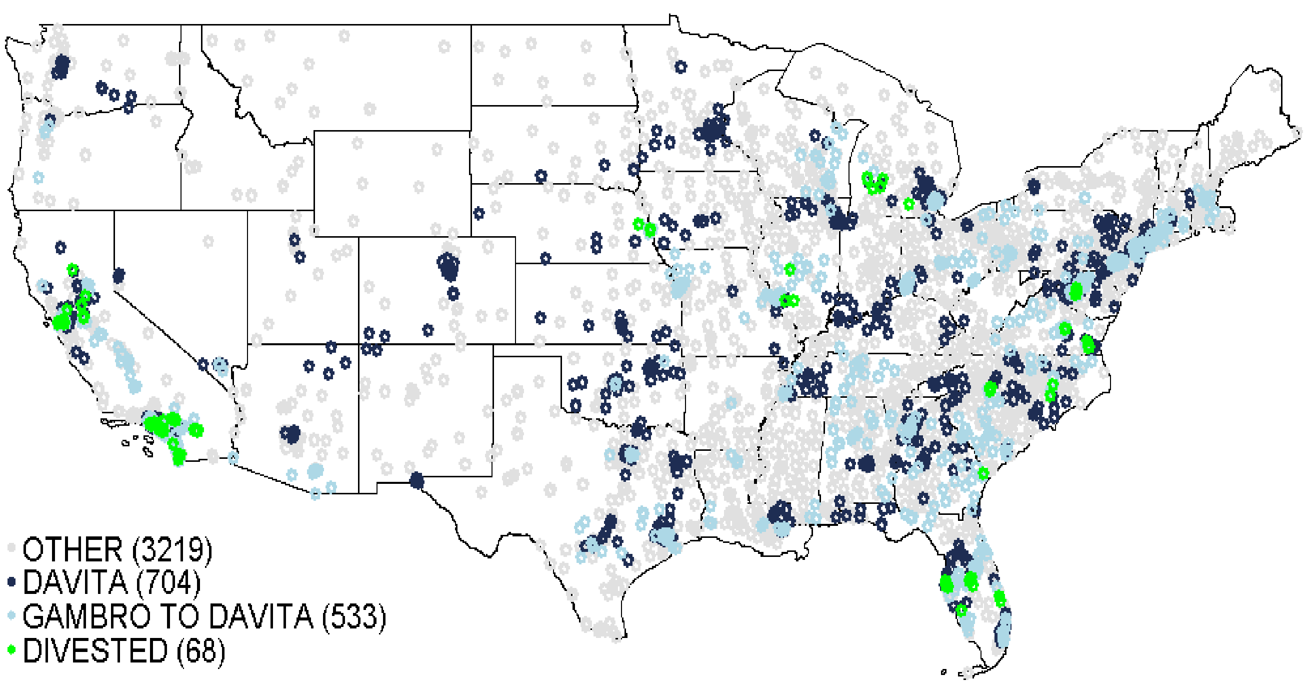
<u>Note</u>: Restricted to HSAs with dialysis facilities in both 2000 and 2009, and not served by a single firm in both years. N=773.

Figure 4: Merging Clinic Locations

Fresenius and RCG

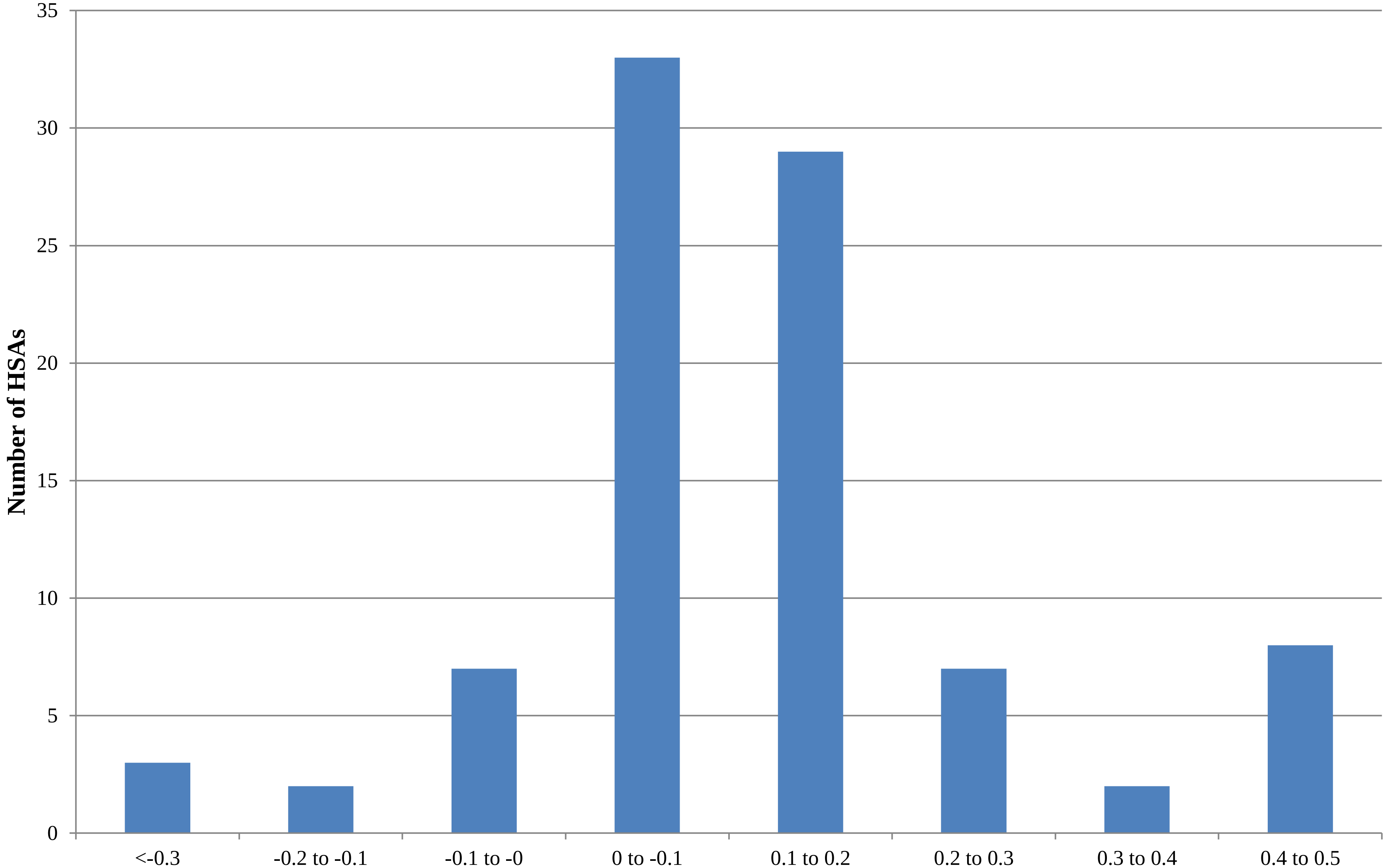


DaVita and Gambro



Note: Legends refer to numbers of clinics in parentheses. Data year is 2005. The sample is the 4,546 clinics used to calculate HHI (ie exclude facilities ever having mostly pediatric patients), 22 of which are in Hawaii or Alaska and therefore omitted from the graphic and facility counts.

Figure 5: Predicted Impact of 2005/2006 Mergers ("Simulated Change in HHI")



Note: Restricted to HSAs with non-zero values of "simulated change in HHI". N=91.

Figure 6: Relationship between Local Market Concentration and Dialysis Quality (Reduced Form Estimates)

Market definition = HSA, Study Period = 2000-2009

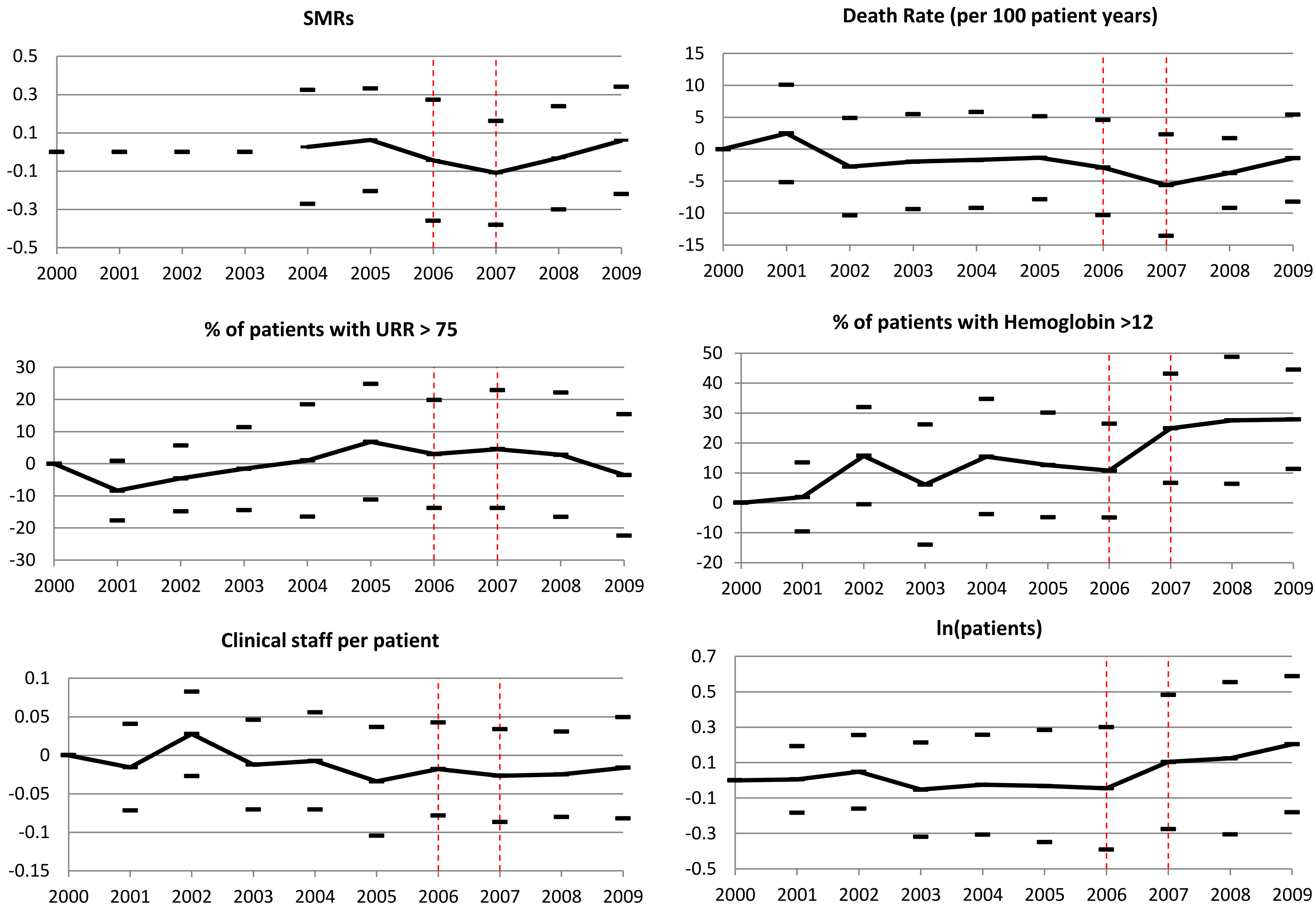
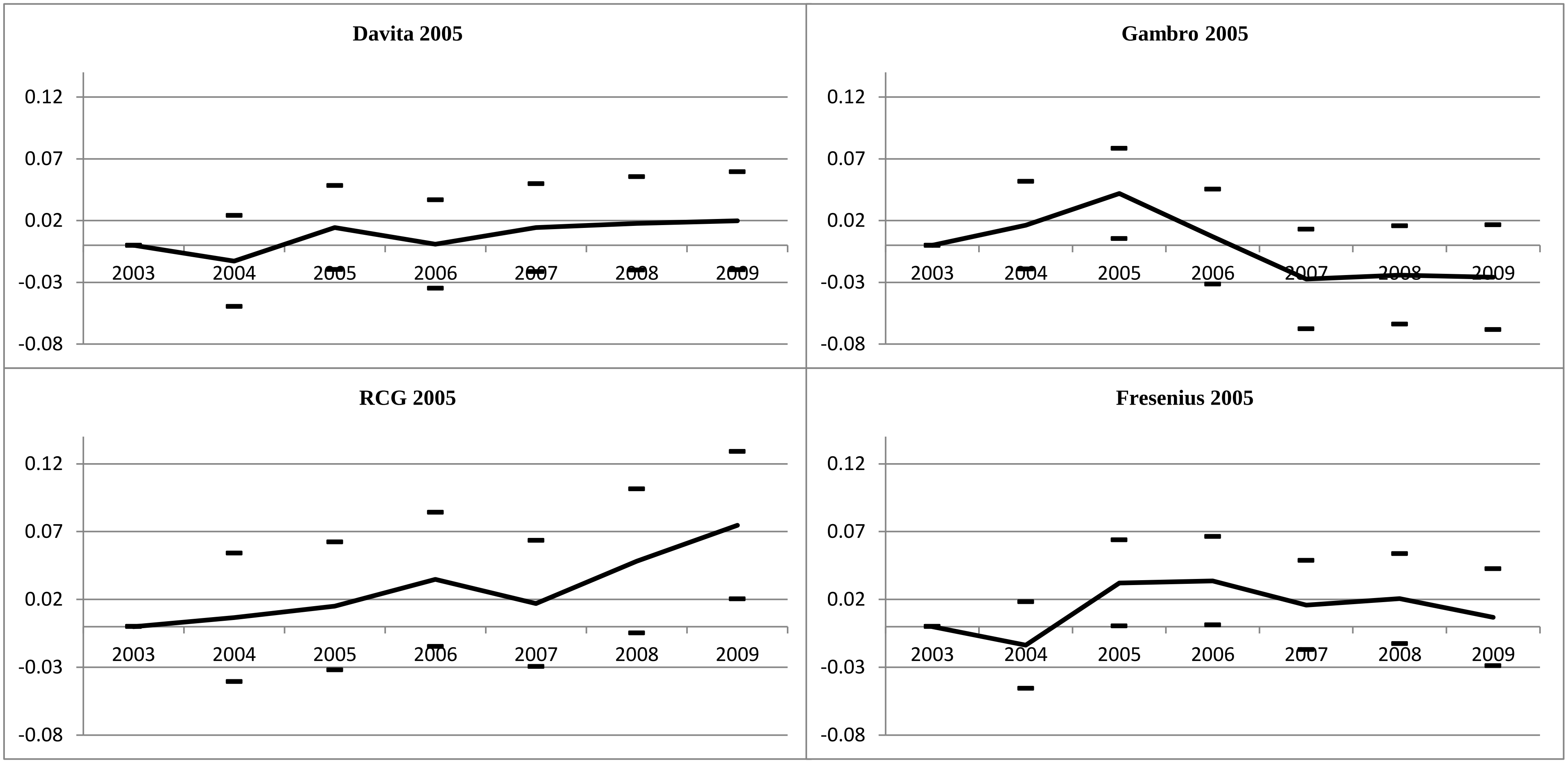


Figure 7: Effect of Mergers on Chain Quality



Notes: Data is at the facility-year level. Sample excludes facilities in markets with >10 hospitals, facilities with >50 percent home-based treatments or >50 percent pediatric patients in any year of the data. All specifications include facility and year fixed effects. Observations are estimated by WLS using facility average end-of-year patient counts as weights. Standard errors are clustered by facility.

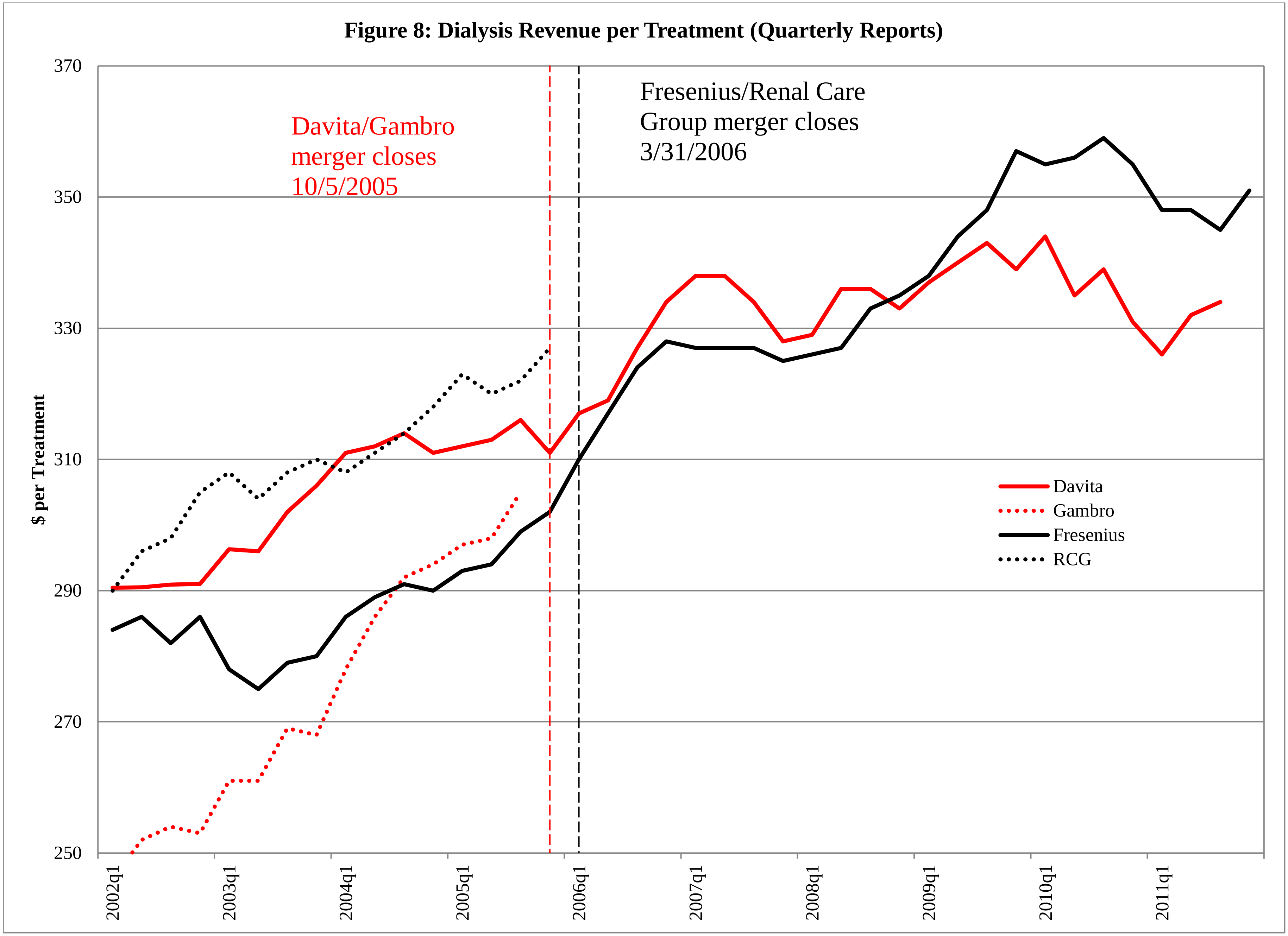


Figure 8: Dialysis Revenue per Treatment (Quarterly Reports)
$ per Treatment
370
350
330
310
290
270
250
Davita/Gambro
merger closes
10/5/2005
Fresenius/Renal Care
Group merger closes
3/31/2006
Davita
Gambro
Fresenius
RCG
2002q1
2003q1
2004q1
2005q1
2006q1
2007q1
2008q1
2009q1
2010q1
2011q1

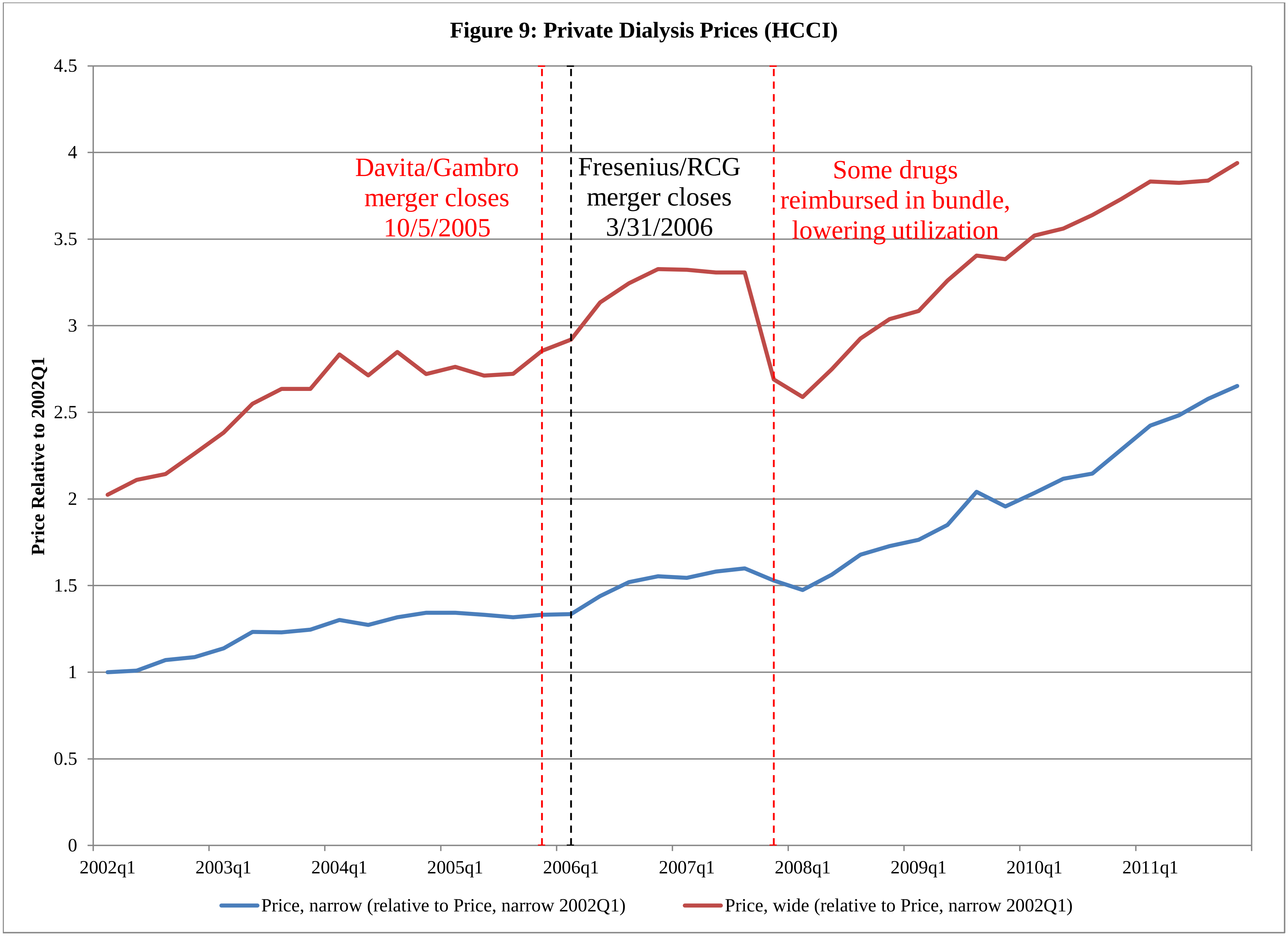
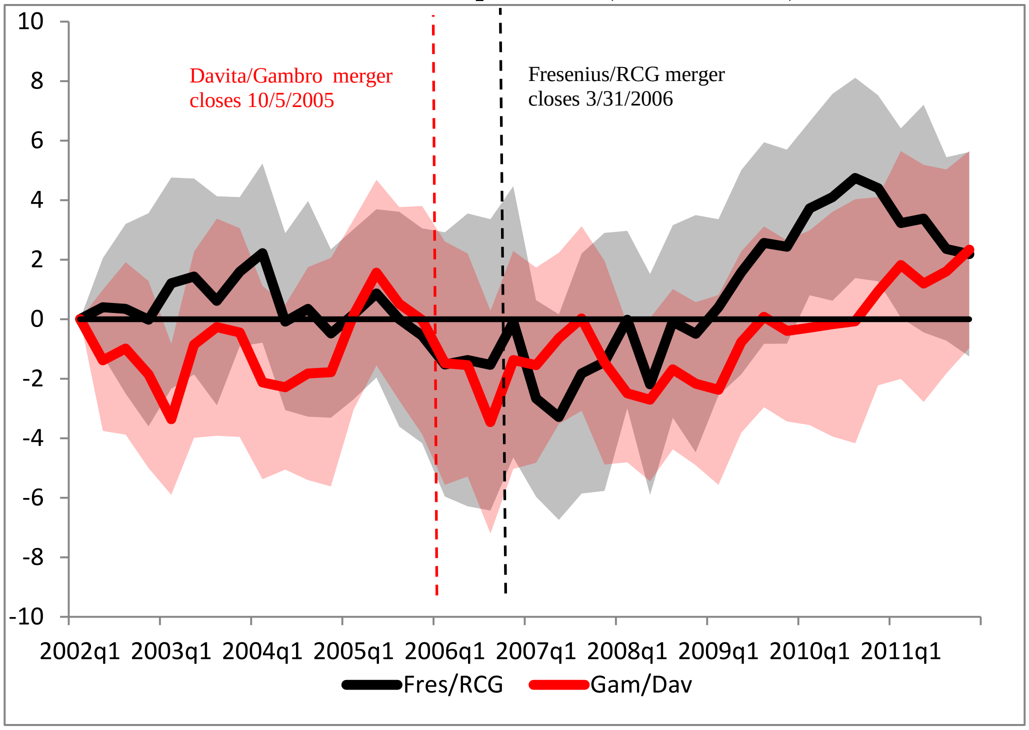


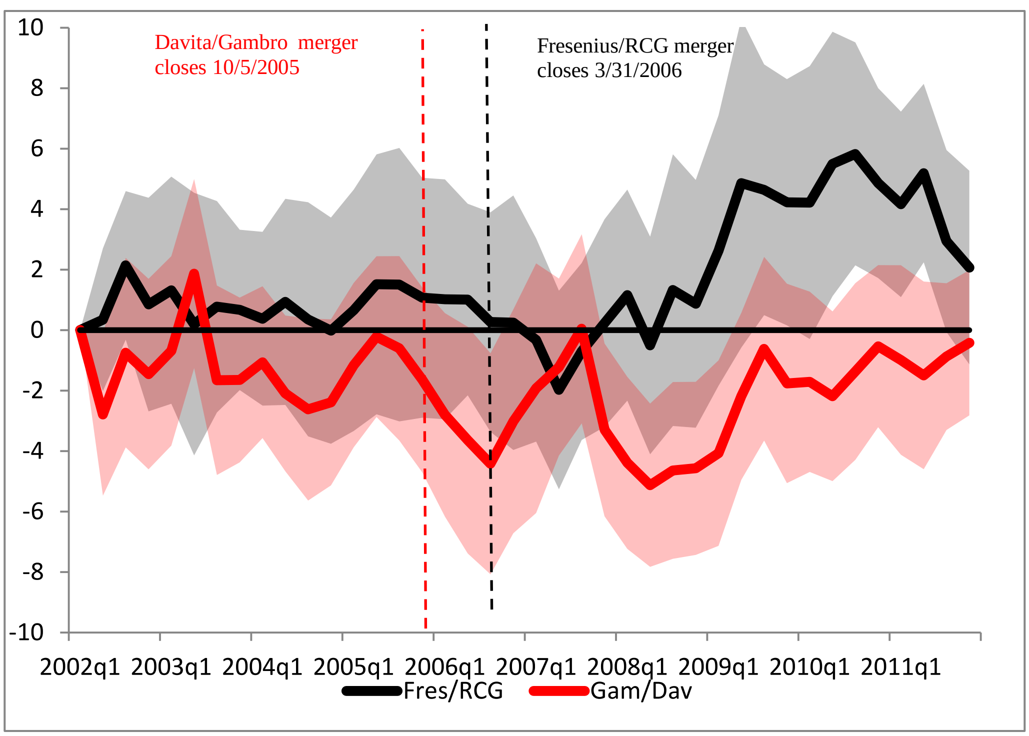
Figure 9: Private Dialysis Prices (HCCI)
Price Relative to 2002Q1
Davita/Gambro merger closes 10/5/2005
Fresenius/RCG merger closes 3/31/2006
Some drugs reimbursed in bundle, lowering utilization
0
0.5
1
1.5
2
2.5
3
3.5
4
4.5
2002q1
2003q1
2004q1
2005q1
2006q1
2007q1
2008q1
2009q1
2010q1
2011q1
Price, narrow (relative to Price, narrow 2002Q1)
Price, wide (relative to Price, narrow 2002Q1)

Figure 10: Reduced Form Relationship between ΔHHI and Prices

Panel A: Dep Var = ln(Price, narrow)



Panel B: Dep Var = ln(Price, wide)



Notes: Controls are Lagged unemployment, enrollee characteristics (<45, 45-54, 55+); gender; % HMO. Regressions are weighted by average HSA treatments from HCCI. Standard errors are clusterred by LEHID market. 95% contidence intervals are presented in shading.

Table 1: Summary Statistics, Dialysis Facilities

Facility-level data	**2000-2009**	**2000**	**2005**	**2009**
Independent Variables				
SMR*	1.034	-	0.994	1.024
	(0.344)		(0.319)	(0.343)
Death rate	22.69	23.75	23.28	20.88
	(8.423)	(8.694)	(8.173)	(7.762)
Clinical staff per patient	0.202	0.219	0.200	0.186
	(0.0507)	(0.0575)	(0.0437)	(0.0395)
Hemoglobin > 12	34.10	27.61	49.84	15.90
	(20.98)	(16.78)	(18.74)	(10.34)
Urea Reduction ratio > 75	42.74	34.14	44.36	51.99
	(18.02)	(16.80)	(16.86)	(16.36)
Controls				
% Medicare	88.40	85.22	89.33	88.37
	(7.741)	(10.20)	(6.543)	(7.109)
% hemodialysis	93.62	94.16	94.00	93.44
	(9.959)	(9.148)	(9.949)	(10.90)
% female	46.43	47.58	46.20	45.15
	(8.596)	(8.777)	(8.562)	(8.423)
% black	36.42	37.28	36.27	35.48
	(30.65)	(30.65)	(30.72)	(30.46)
Avg yrs on ESRD	3.978	3.712	4.023	4.325
	(0.985)	(0.940)	(0.965)	(1.002)
Number of observations**	51,598	3,701	4,546	5,276
Clinic Characteristics				
% for-profit	0.909	0.912	0.905	0.913
	(0.287)	(0.283)	(0.294)	(0.282)
% in chains of 100+ clinics	0.879	0.876	0.884	0.886
	(0.326)	(0.330)	(0.321)	(0.318)
Market-level data	**2000-2009**	**2000**	**2005**	**2009**
HHI	8471.3	8603.0	8437.2	8334.0
	(2386.7)	(2349.8)	(2419.1)	(2417.7)
Number of firms	1.599	1.582	1.613	1.625
	(1.471)	(1.593)	(1.497)	(1.360)
Number of clinics	2.262	2.067	2.269	2.448
	(3.679)	(3.246)	(3.690)	(4.106)
Number of HSAs	-	1,795	2,007	2,160

Notes: All statistics are unweighted. The unit of observation is the facility-year. The sample is all observations used in calculating HHIs, and therefore excludes facilities with >50% pediatric patients in any year.

* Data for SMRs is only available from 2003-2009

**Number of observations differs across variables. Regression results include counts for each outcome.

Table 2: Summary Statistics, Price Analysis

	Whole sample				
	2003	**2004**	**2005**	**2006**	**2007**
Lagged dialysis HHI	0.20	0.26	0.27	0.29	0.34
	(0.11)	(0.14)	(0.14)	(0.14)	(0.1)
Lagged insurance HHI	0.22	0.26	0.29	0.31	0.29
	(0.07)	(0.08)	(0.09)	(0.13)	(0.08)
Price per treatment	375	402	469	470	492
	(284)	(248)	(312)	(305)	(253)
Ancillaries price per treatment	391	395	433	438	409
	(1134)	(332)	(351)	(414)	(362)
Lagged Average Annual Per Capita	6219	6493	6968	7531	8195
Cost per Medicare enrollee	(847)	(815)	(869)	(991)	(1090)
Unique markets	94	122	128	127	137
Unique employers	18	45	62	62	104
Number of markets x employers	183	544	736	710	1335

	Affected markets				
	2003	**2004**	**2005**	**2006**	**2007**
Lagged dialysis HHI	0.18	0.21	0.23	0.24	0.32
	(0.08)	(0.09)	(0.09)	(0.08)	(0.08)
Lagged insurance HHI	0.22	0.25	0.27	0.29	0.28
	(0.07)	(0.07)	(0.08)	(0.1)	(0.07)
Price per treatment	403	423	497	493	498
	(228)	(249)	(317)	(321)	(249)
Ancillaries price per treatment	366	399	450	451	401
	(745)	(347)	(372)	(453)	(358)
Lagged Average Annual Per Capita	6219	6542	7055	7701	8353
Cost per Medicare enrollee	(899)	(873)	(923)	(1048)	(1084)
Unique markets	68	81	84	85	85
Unique employers	18	44	62	61	104
Number of markets x employers	144	434	599	598	1055

Notes: All statistics weighted by the number of treatments. The unit of observation is the LEHID market-employer-year.

Table 3: Major Acquisitions in Dialysis Industry, 2000-2011

Acquirer	Target	Merger announcment and completion dates	# Target Facilities	# Facilities divested
Fresenius	Everest Healthcare Service Corporation	Announced: 11/2/2000 Completed: 1/ 9/2001	74	0
Renal Care Group	National Nephrology Association	Announced: 2/2/2004 Completed: 4/2/2004	87	0
Davita	Gambro	Announced: 12/07/04 Completed: 10/5/2005	565	68 (to Renal Advantage)
Fresenius	Renal Care Group	Announced: 5/4/2005 Completed: 3/31/2006	>450	103 (to DSI)
Renal Advantage, Inc	Liberty Dialysis	Announced: 11/4/2010 Completed: 12/22/2010	112	0
Davita	DSI	Announced: 2/04/2011 Completed: 9/02/2011	106	29 (to Dialysis, Newco)
Fresenius	Renal Advantage	Announced: 8/02/2011 Completed: 4/02/2012	260	60 (to Dialysis, Newco)

Notes: Includes all mergers and acquisitions where the target has at least 70 clinics. Chain sizes are as reported in press releases and therefore do not correspond exactly to numbers from our data.

Table 4: Relationship between Local Market Concentration and Dialysis Quality

(Data is for 2000-2009, except where otherwise noted)

	Dep Var = SMR (2003-2009) (mean=.99, std dev=.33)			Dep Var = Observed death rate (mean=23, std dev=9)			Dep Var = % of patients with URR> 75 (mean=44, std dev=18)		
	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
Lagged HHI	-0.013	N/A	-0.012	-0.339	N/A	-0.247	0.096	-0.627	-1.193
	(0.026)	N/A	(0.026)	(0.461)	N/A	(0.463)	(1.458)	(1.537)	(1.352)
Demographic controls (p-value from F-test)								0.000	0.000
Individual chain*year interactions (p-value from F-test)			0.000			0.000			0.000
Adjusted R2	0.021	N/A	0.023	0.037	N/A	0.040	0.216	0.221	0.319
# of Observations	27,044	N/A	27,044	36,717	N/A	36,717	35,898	33,688	33,688

	Dep Var = % of patients with hemoglobin >12 (mean=36, std dev=21)			Dep Var = clinical staff per patient (mean=.2, std dev=.05)		
	(1)	(2)	(3)	(1)	(2)	(3)
Lagged HHI	0.451	0.200	-0.531	0.001	0.001	-0.004
	(1.687)	(1.813)	(1.592)	(0.004)	(0.004)	(0.004)
Demographic controls (p-value from F-test)		0.000	0.000		0.000	0.001
Individual chain*year interactions (p-value from F-test)			0.000			0.000
Adjusted R2	0.427	0.433	0.512	0.076	0.085	0.193
# of Observations	36,211	33,960	33,960	31,255	29,211	29,211

Notes: Data is at the facility-year level. Sample excludes facilities in markets with >10 hospitals, facilities with >50 percent home-based treatments or >50 percent pediatric patients in any year of the data. HHI is scaled from 0 to 1. Demographic controls include lags of: % medicare, % hemodialysis, gender, average years on dialysis. Individual chain*year interactions are included for chains with >70 clinics in that year. All specifications include facility and year fixed effects. Observations are estimated by WLS using facility average end of year patient counts as weights. Standard errors are clustered by HSA.

*** p<.001, ** p<0.01, * p<0.05

Table 5: Effect of the 2005/2006 Mergers on Market Concentration

	Dependent variable = HHI
(Sim ΔHHI)* (Year==2000)	omitted
(Sim ΔHHI)* (Year==2001)	-0.195
	(0.135)
(Sim ΔHHI)* (Year==2002)	-0.160
	(0.151)
(Sim ΔHHI)* (Year==2003)	-0.115
	(0.161)
(Sim ΔHHI)* (Year==2004)	-0.117
	(0.169)
(Sim ΔHHI)* (Year==2005)	-0.126
	(0.167)
(Sim ΔHHI)* (Year==2006)	0.758***
	(0.186)
(Sim ΔHHI)* (Year==2007)	0.651***
	(0.188)
(Sim ΔHHI)* (Year==2008)	0.636**
	(0.189)
(Sim ΔHHI)* (Year==2009)	0.649**
	(0.190)
Market Definition	HSA
Number of observations	720
Adjusted R2	0.826

Notes: The unit of observation is the market-year, where markets are HSAs. Excludes markets with >10 hospitals, markets with no merger induced change in HHI and markets with large negative merger induced decreases in HHIs. All specifications include market and year fixed effects. Observations are estimated by WLS using market end-of-year patient counts as weights. Standard errors are clustered by market.
*** p<.001, ** p<0.01, * p<0.05

Table 6: Relationship between Merger-Induced Changes in Local Market Concentration and Dialysis Quality (Reduced Form Estimates)

(Data is for 2000-2009, except where otherwise noted.)

	Dep Var = SMR (2003-2009) (mean=1.01, std dev=.33)				Dep Var = Observed death rate (mean=22, std dev=9)				Dep Var = % of patients with URR> 75 (mean=42, std dev=17)			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
(Sim ΔHHI)*(Year>2006)	-0.123	N/A	-0.116	-0.103	-2.002	N/A	-1.830	-1.481	-0.659	-1.649	-1.614	-4.279
	(0.153)	N/A	(0.151)	(0.166)	(2.673)	N/A	(2.551)	(2.675)	(7.569)	(7.686)	(7.467)	(6.984)
Market specific trends	0.000	N/A	0.000	0.000	0.000	N/A	0.000	0.000	0.000	0.000	0.000	0.000
Demographic controls		N/A				N/A				0.002	0.001	0.001
Target*post interactions			0.069	0.010			0.035	0.036			0.218	0.079
Individual chain* year interactions				0.000				0.000				0.000
R-squared	0.031	N/A	0.033	0.036	0.085	N/A	0.087	0.090	0.291	0.302	0.305	0.389
# of Observations	3,729	N/A	3,729	3,729	5,096	N/A	5,096	5,096	4,971	4,632	4,632	4,632

	Dep Var = % of patients with hemoglobin >12 (mean=37, std dev=20)				Dep Var = clinical staff per patient (mean=.2, std dev=.05)			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
(Sim ΔHHI)*(Year>2006)	11.063	9.613	8.648	9.798	-0.006	-0.013	-0.017	0.006
	(9.405)	(8.999)	(8.961)	(7.729)	(0.019)	(0.019)	(0.018)	(0.020)
Market specific trends	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Demographic controls		0.038	0.063	0.192		0.092	0.134	0.039
Target*post interactions			0.000	0.013			0.000	0.287
Individual chain* year interactions				0.000				0.000
R-squared	0.500	0.512	0.518	0.586	0.108	0.121	0.143	0.246
# of Observations	5,039	4,692	4,692	4,692	4,641	4,313	4,313	4,313

Notes: Data is at the facility-year level. Standard errors are in parentheses. Sample excludes facilities in markets with >10 hospitals, facilities in markets with no merger induced change in HHI, facilities with large divestiture-induced decreases in HHIs, facilities with >50 percent home-based treatments or >50 percent pediatric patients in any year of the data. HHI is scaled from 0 to 1. Demographic controls include lags of: % Medicare, % hemodialysis, gender, average years on dialysis. Target*post interactions are as specified in the text. Individual chain*year interactions are included for chains with >70 clinics in that year. All specifications include facility and year fixed effects. Observations are estimated by WLS using facility average end-of-year patient counts as weights. Standard errors are clustered by HSA.

*** p<0.01, ** p<0.05, * p<0.1

Table 7: Relationship between Local Market Concentration and Dialysis Quality (Reduced Form Estimates)

Market definition = HSA, Study Period = 2000-2009

	Dep Var = SMR (2003-2009) (mean=1.01, std dev=.33)				Dep Var = Observed death rate (mean=22, std dev=9)				Dep Var = % of patients with URR> 75+ (mean=42, std dev=17)			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
(Sim ΔHHI)* (Year>2006)* competitor	-0.146	N/A	-0.251	-0.130	-0.165	N/A	-2.387	-0.429	1.614	-0.600	2.147	-2.238
	(0.275)	N/A	(0.286)	(0.313)	(4.707)	N/A	(4.919)	(5.439)	(17.186)	(18.111)	(18.831)	(17.887)
(Sim ΔHHI)* (Year>2006)* merging	-0.116	N/A	-0.077	-0.094	-2.465	N/A	-1.684	-1.775	-1.210	-1.850	-2.373	-4.722
	(0.143)	N/A	(0.140)	(0.155)	(2.604)	N/A	(2.387)	(2.555)	(6.696)	(6.979)	(6.548)	(6.764)
Market specific trends	0.000	N/A	0.000	0.000	0.000	N/A	0.000	0.000	0.000	0.000	0.000	0.000
Demographic controls		N/A				N/A				0.002	0.001	0.001
Target*post interactions			0.059	0.011			0.041	0.034			0.222	0.076
Individual chain*year interactions				0.000				0.000				0.000
R-squared	0.031	N/A	0.033	0.036	0.085	N/A	0.087	0.089	0.291	0.302	0.305	0.389
# of Observations	3,729	N/A	3,729	3,729	5,096	N/A	5,096	5,096	4,971	4,632	4,632	4,632

	Dep Var = % of patients with Hemoglobin >12 (mean=37, std dev=20)				Dep Var = clinical staff per patient (mean=.2, std dev=.05)				Dep Var = ln(patients) (mean=4.2, std dev=.6)			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
(Sim ΔHHI)* (Year>2006)* competitor	14.361	9.288	24.462*	0.061	-0.062*	-0.095**	-0.053	0.060*	0.335**	0.195	0.084	-0.065
	(13.349)	(14.306)	(13.781)	(14.507)	(0.037)	(0.042)	(0.038)	(0.034)	(0.165)	(0.178)	(0.163)	(0.193)
(Sim ΔHHI)* (Year>2006)* merging	10.257	9.676	5.443	11.896	0.006	0.000	-0.011	-0.004	0.081	0.121	0.153*	0.218**
	(9.976)	(9.360)	(9.668)	(7.721)	(0.024)	(0.023)	(0.020)	(0.020)	(0.102)	(0.091)	(0.092)	(0.092)
Market specific trends	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Demographic controls		0.039	0.057	0.205		0.102	0.129	0.037		0.002	0.002	0.000
Target*post interactions			0.000	0.014			0.000	0.300			0.054	0.028
Individual chain*year interactions				0.000				0.000				0.000
R-squared	0.500	0.512	0.519	0.586	0.110	0.124	0.144	0.247	0.090	0.108	0.111	0.118
# of Observations	5,039	4,692	4,692	4,692	4,641	4,313	4,313	4,313	5,117	4,717	4,717	4,717

Notes: Data is at the facility-year level. Sample excludes facilities in markets with >10 hospitals, facilities in markets with no merger induced change in HHI, facilities with large merge induced decreases in HHIs, facilities with >50 percent home-based treatments or >50 percent pediatric patients in any year of the data. HHI is scaled from 0 to 1. Demographic controls include lags of: % medicare, % hemodialysis, gender, average years on dialysis. Target*post interactions are as specified in the text. Individual chain*year interactions are included for chains with >70 clinics in that year. All specifications include facility and year fixed effects. Observations are estimated by WLS using facility average end of year patient counts as weights. Standard errors are clustered by HSA.
*** p<0.01, ** p<0.05, * p<0.1

Table 8: Effect of Merging Firm's Combined Shares on Quality

	Dep Var = SMR (2003-2009)
(HSA % merging in 2005)*(year==2003)	omitted
(HSA % merging in 2005)*(year==2004)	-0.007
	(0.015)
(HSA % merging in 2005)*(year==2005)	0.031
	(0.015)
(HSA % merging in 2005)*(year==2006)	0.025
	(0.016)
(HSA % merging in 2005)*(year==2007)	0.004
	(0.016)
(HSA % merging in 2005)*(year==2008)	0.022
	(0.017)
(HSA % merging in 2005)*(year==2009)	0.018
	(0.018)

Notes: Data is at the facility-year level. Sample excludes facilities in markets with >10 hospitals, facilities with >50 percent home-based treatments or >50 percent pediatric patients in any year of the data. All specifications include facility and year fixed effects. Observations are estimated by WLS using facility average end-of-year patient counts as weights. Standard errors are clustered by HSA.

Table 9: Effect of the 2005/2006 Mergers on Prices

Panel A: Regression Coefficients

	Dep Var =ln(Price, narrow)			Dep Var =ln(Price, wide)		
	(1)	(2)	(3)	(4)	(5)	(6)
Merger effects						
$\Delta HHI_{Dav/Gam}$*(year>2006)	0.884	-1.186		-0.073	-1.824	
	(1.296)	(1.000)		(1.045)	(1.133)	
$\Delta HHI_{Fres/RCG}$*(year>2006)	-0.111	-2.921***		0.369	-1.787	
	(1.136)	(0.845)		(1.204)	(1.307)	
$\Delta HHI_{Combined}$*(year>2006)			-1.699**			-1.806**
			(0.759)			(0.872)
Interactions with linear post-merger trend						
$\Delta HHI_{Dav/Gam}$*(year>2006)*quarter		0.217***			0.184**	
		(0.082)			(0.077)	
$\Delta HHI_{Fres/RCG}$*(year>2006)*quarter		0.294***			0.226***	
		(0.089)			(0.066)	
$\Delta HHI_{Combined}$*(year>2006)*quarter			0.239***			0.196***
			(0.064)			(0.060)
2005 shares of each merging chain* (year>2006)*quarter		Y	Y		Y	Y
2005 shares of each merging chain* (year>2006)	Y	Y	Y	Y	Y	Y
R-sq	0.242	0.247	0.247	0.091	0.099	0.099
Number of Observations	26634	26634	26634	26742	26742	26742

Panel B: Implied Coefficients for 2011Q4

	Dep Var =ln(Price, narrow)			Dep Var =ln(Price, wide)		
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta HHI_{Dav/Gam}$	N/A	2.932		N/A	1.665	
		(1.885)			(1.400)	
$\Delta HHI_{Fres/RCG}$	N/A	2.672		N/A	2.515*	
		(1.807)			(1.410)	
$\Delta HHI_{Combined}$			2.842*			1.914*
			(1.460)			(1.134)

*Notes: All specifications include Lagged unemployment and enrollee characteristics. Linear trends are labeled "quarter," and begin in 2007Q1 and increment by one each quarter. Regressions are weighted by average treatments for each HSA. SEs in () are clustered by LEHID market. * p<0.10, ** p<0.05, *** p<.01*

Appendix Table 1: First Stages

(Data is for 2000-2009)

	Dep Var = HHI			
	(1)	(2)	(3)	(4)
(Sim ΔHHI)* (Year==2000)	omitted	omitted	omitted	omitted
(Sim ΔHHI)* (Year==2001)	omitted	omitted	omitted	omitted
(Sim ΔHHI)* (Year==2002)	-0.00137	0.0135	0.0135	0.0123
	(-0.02)	(0.19)	(0.19)	(0.18)
(Sim ΔHHI)* (Year==2003)	0.0735	0.0971	0.0972	0.0932
	(0.67)	(0.89)	(0.89)	(0.93)
(Sim ΔHHI)* (Year==2004)	0.0996	0.146	0.146	0.139
	(0.68)	(1.01)	(1.01)	(1.05)
(Sim ΔHHI)* (Year==2005)	0.198	0.226	0.226	0.217
	(1.07)	(1.19)	(1.19)	(1.25)
(Sim ΔHHI)* (Year==2006)	1.186***	1.223***	1.223***	1.201***
	(5.25)	(5.33)	(5.33)	(5.72)
(Sim ΔHHI)* (Year==2007)	1.176***	1.212***	1.213***	1.188***
	(4.27)	(4.28)	(4.29)	(4.53)
(Sim ΔHHI)* (Year==2008)	1.190***	1.233***	1.234***	1.217***
	(3.78)	(3.81)	(3.81)	(4.04)
(Sim ΔHHI)* (Year==2009)	1.256***	1.299***	1.300***	1.290***
	(3.62)	(3.62)	(3.62)	(3.87)
Market specific trends	0.000	0.000	0.000	0.000
Demographic controls		0.272	0.285	0.601
Target*post interactions			0.658	0.803
Individual chain*year interactions				0.000
R-squared	0.803	0.808	0.808	0.815
# of Observations	5,117	4,717	4,717	4,717

Notes: Data is at the facility-year level. Standard errors are in parentheses. Sample excludes facilities in markets with >10 hospitals, facilities in markets with no merger induced change in HHI, facilities with large divestiture-induced decreases in HHIs, facilities with >50 percent home-based treatments or >50 percent pediatric patients in any year of the data. HHI is scaled from 0 to 1. Demographic controls include lags of: % Medicare, % hemodialysis, gender, average years on dialysis. Target*post interactions are as specified in the text. Individual chain*year interactions are included for chains with >70 clinics in that year. All specifications include facility and year fixed effects. Observations are estimated by WLS using facility average end-of-year patient counts as weights. Standard errors are clustered by HSA.

*** p<0.01, ** p<0.05, * p<0.1