# Waiting for Primary Care and Health Outcomes among Older Patients With Diabetes

Julia C. Prentice, PhD

B. Graeme Fincke, MD

Donald R. Miller, ScD

Steven D. Pizer, PhD

**Acknowledgements**
Funding for this research was provided by Grant No: IAD-06-112 and IIR 04-233from the Health Services Research and Development Service, Department of Veterans Affairs and Grant No: 62967 from the Health Care Financing and Organization Initiative under the Robert Wood Johnson Foundation. The authors are indebted to Matthew Neuman and John Gardner, PhD for programming support.

**Disclaimer**
The views expressed in this article are those of the authors and do not necessarily represent the position or policy of the Department of Veteran Affairs or Boston University.

**Abstract**

**Objective:** To measure the relationship between days spent waiting for primary care and health outcomes among patients diagnosed with diabetes, especially among the elderly population.

**Data Source:** Secondary data from VA administrative databases and Medicare claims.

**Study Design:** This is a retrospective observational study. Outcome variables include primary care utilization, mortality, heart attack, stroke and ambulatory-care sensitive condition (ACSC) hospitalization. The main explanatory variable of interest is VA primary care wait time. Negative binomial models predict utilization and stacked logistic regression models predict the probability of experiencing each health outcome. Models are stratified by the presence of a selected health condition and age.

**Principal Findings:** Longer wait times were predicted to decrease utilization between 2 and 4%. There was no significant relationship between wait times and health outcomes for the overall sample. In stratified analyses, longer waits were associated with undesirable outcomes for those over age 70 with one of the selected health conditions or in certain narrower five-year age groups, but the overall pattern of results does not indicate a systematic and significant effect.

**Conclusions:** There was a modest effect of long wait times on primary care utilization but no robust effect of longer wait times on health outcomes. Waiting for care did not significantly compromise long-term health outcomes for veterans with diabetes.

**Key Words:** Access to care, mortality, primary care wait times, ambulatory care-sensitive conditions, diabetes

**Background**

Delayed access to health care in the United States frustrates patients and concerns policymakers (Cunningham and Felland 2008; Lukas et al. 2004; U.S. GAO 1993). The Institute of Medicine's Committee on the Quality of Health Care in America (2001) highlighted reducing delays as one of six aims for improving the quality of America's health care system. Beyond the inconvenience to patients of not being able to access physician care when wanted, policymakers assume delayed access to health care negatively affects health outcomes through decreased or foregone health care utilization that delays diagnosis and treatment (IOM 2001; Kenagy, Berwick, and Shore 1999). Survey data support this link. In the 2007 Community Tracking Study, 35% of respondents with foregone or delayed health care did so because they could not get an appointment soon enough (Cunningham and Felland 2008).

Concern about waiting times is particularly timely now because major health reform that expands access to insurance coverage recently became law. The expansion of health insurance without a concomitant expansion in the supply of health care services to respond to the increased demand will result in longer waits (Buchmueller et al. 2005; CBO 2008).

Due to limited data on wait times, research testing the relationship between waiting for care and health outcomes is rare. One of the only sources of administrative data on wait times is the Veterans Health Administration (VA). Two separate studies using VA data support concerns about delayed access leading to poor health outcomes among a sample of veterans receiving geriatric clinic care. Prentice and Pizer (2007) found that elderly veterans who visited VA facilities with an average wait time of 31 days or more were significantly more likely to die compared to veterans who visited VA facilities with shorter waits. This study used random effects to account for unobservable characteristics of facilities, such as facility quality. If these characteristics were correlated with wait times, the estimated effect was confounded. Using the same sample, a subsequent study examined the relationship between waiting for outpatient care and the risk of ambulatory-care sensitive condition (ACSC) hospitalization (Prentice and Pizer 2008). ACSC hospitalizations are hospitalizations that are potentially avoidable with timely, high-quality outpatient care (AHRQ 2001, AHRQ 2010). Geriatric clinic patients who visited VA facilities with wait times of 29 days or more had a significantly higher probability of having an ACSC hospitalization. This study included facility-level fixed effects, which controlled for facility quality characteristics that are constant over time. Taken together, these studies suggested that delayed access to care was problematic in an elderly population with a wide range of clinical conditions. The present study expands our understanding of the effect of wait times on health care utilization and outcomes in a clinically homogeneous population by focusing on a population diagnosed with diabetes.

*Managing Diabetes in an Outpatient Setting*

Diabetes is a chronic condition that relies on timely outpatient care for appropriate management. Complications from diabetes are minimized when providers focus on controlling blood glucose ($HbA_{1c}$) levels and hypertension and monitor early signs of potential complications such as heart attack or foot ulcers (ADA 2005; Ross 2004; Stratton et al. 2000; DCCT 1996, 2000). Consequently, diabetes is a major focus of the ACSC hospitalizations (AHRQ 2001; AHRQ 2010).

Elderly patients with diabetes may be at especially high risk of experiencing poor health outcomes if faced with delays in outpatient care. For example, as age increases, the physiological mechanisms that regulate blood sugar are impaired, increasing the risk of hypoglycemia (Chau and Edelman 2001; Hornick and Aron 2008). Geriatricians have defined frailty as a biologic syndrome of decreased physiological reserves that increase vulnerability to stressors. Among frail patients, delays in responding to small changes in health status may significantly increase the risk of more serious health outcomes (Fried et al. 2004). A higher prevalence of complex comorbid illnesses may also complicate diabetes management (Durso 2006). Finally, many elderly patients have been living with diabetes for a long period of time.  This sustained metabolic imbalance increases the risk and severity of microvascular or macrovascular complications compared to newly diagnosed patients (Durso 2006; Fowler 2008).

## Methods

### Study Period and Study Population

Data are from VA administrative records and Medicare claims.  This study measures outcomes from January 1, 2002 through December 31, 2003 with calendar year 2001 used as a baseline year for risk adjustment.  We extract the sample from the VA Diabetes Epidemiology Cohorts database, a registry of all VA patients with diabetes since 1998.  The criteria for identifying diabetes includes a prescription for diabetes medication in the current year and/or 2+ visits or stays with a diabetes code for inpatient and/or outpatient visits (VA and Medicare) over a two-month period (Miller, Safford, and Pogach 2004).  All veterans diagnosed with diabetes before 2001 are eligible to enter the sample. Separating the sample selection period (diabetes diagnosis before 2001) from the measurement of outcomes (2002 and 2003) minimizes potential sample selection bias (Prentice and Pizer 2007, 2008).

In order to maximize the completeness of utilization and diagnosis records we further limit the sample to veterans who were also eligible for Medicare during 2002 or 2003, according to the Medicare Denominator file.  We exclude individuals who were enrolled in a Medicare HMO in 2001 because claims data for these enrollees are not available.

Eligibility for Medicare increases our confidence that we can observe all utilization among our sample of veterans between the VA and Medicare data.  We also want to maximize our chances of observing a relationship between waiting times and health outcomes. Veterans who rely heavily on Medicare are less likely to be affected by VA waiting times.  Consequently, we further restrict the sample to veterans with no Medicare outpatient visits in 2001.

Waiting Time Data

The main explanatory variable of interest is the wait until the next available appointment at a VA facility.[1] Services in the VA are provided at a parent station, such as a medical center, or a sub-station, such as a community-based outpatient clinic. Wait times are kept for each type of clinic appointment (e.g. urology) within a parent station and for all clinic appointments at a sub-station. To create a performance measure for tracking wait times, the VA aggregates wait times by appointment type to the parent station level on a monthly basis using a weighted average. The weight is the number of next available appointments at each sub-station that operate under a VA parent station (Baar 2005a, 2005b).

The wait time measure in this study focuses on clinics that provide primary care. Primary care doctors in the VA manage diabetes and refer patients to specialty care as needed (Helmer et al. 2008). Consequently, primary care waits are a measure of the initial delays in accessing the VA system. Our wait time measure focuses on patients who are new to an appointment type (patients who have not had that type of appointment in the last 24 months). This is an overall measure of congestion at the facility. We rely on new patients because VA operations staff found that waiting time data for established patients are unreliable. For these patients, VA scheduling clerks have to distinguish between requests for follow-up appointments that are desired later (e.g. three months out) and urgent care appointments that are desired as soon as possible. Including follow-up appointments in a next available appointment measure may erroneously lengthen the result (Baar2005a, Baar 2005b). VA managers have confirmed that nearly all new patients want the next available appointment (VHA 2008), so focusing on these patients minimizes this concern about data reliability.

Exogenous Wait Time Measure

Previous studies demonstrated that wait times based on services an individual actually uses were inappropriate for studying the effect of wait times on outcomes. Unobserved individual health status affects individual wait times as well as outcomes, because sicker patients were triaged to receive care faster (Prentice and Pizer 2007, 2008). Although statistical controls for observable differences in health status will reduce the severity of this problem, it is not possible to measure health status precisely enough to eliminate it.

To properly isolate the effect of waiting on outcomes, we calculate a wait time that is exogenous to the individual. A detailed example of how this measure is constructed is presented in Appendix A. Briefly, we construct a wait time that would apply to the same "representative" patient at each VA facility. We compute the proportion of each type of primary care clinic appointment used by the entire sample in the baseline period. This proportion is multiplied by the wait in days for each type of primary care appointment at a parent station in a month, and these products are summed for all primary care appointment types in the parent station. Multiplying the wait time for an appointment type by its proportion in the whole sample gives greater weight to the wait times of appointment types that are used more frequently by the entire sample. The result is a facility-level average wait time that is similar to measures used in published studies relating wait times to health outcome (Prentice and Pizer 2007, 2008).

This exogenous wait time is independent of whether or not a specific individual uses a specific VA facility. We take each individual's ZIP code from the Medicare denominator file in 2002 and 2003 and use it to assign them to the nearest VA facility. All individuals who are closest to the same VA facility are assigned the same exogenous facility-level monthly wait time.

Wait time is measured monthly, but it is often longer than 30 days. If a patient requests an appointment and has to wait, the impact of the delay will not typically be felt until the next month when the patient either has the appointment (short wait) or is still waiting (long wait). For this reason we use a lagged wait time measure in our models. It is also possible that the effect of delayed access might take some time to develop into an observable health outcome. To allow for this, we average together the wait times for the previous three months to predict health outcomes in the current month (e.g. average wait for January-March 2002 to predict mortality in April 2002). This measure balances the need to accommodate delays in the development of the outcomes with the need to preserve the variation in wait times found in the data (longer averages smooth variation over time). The lagged 3-month average of wait time changes each month. For ease of interpretation, we scale the wait time measure by 10 days so that the unit of measure for VA wait time is an increase of 10 days.

Facility and Month Fixed Effects and Time Trend

The goal of this paper is to isolate the effect of primary care wait times on primary care utilization and health outcomes. To achieve this goal, a key aspect of our study design is the inclusion of fixed effects for each facility in the models. Including facility fixed effects has two main advantages: 1) alleviating concerns about casemix differences across facilities and 2) controlling for time invariant differences in facility quality.

The facility fixed effect adjusts for between-facility variation in wait times (e.g. wait time differences between Boston VA and Denver VA) and focuses the analysis on the monthly variation in wait times within facilities (e.g. differences between Jan. 02 and Feb. 02 at Boston VA). Since the identifying variation is within facilities and the sample is fixed in the baseline period, individual patients serve as their own controls. We are comparing individual patient outcomes in one month to outcomes for the same patients in other months. This design alleviates concerns about casemix differences across facilities.

The second advantage of including facility fixed effects is that they control for all aspects of facility quality that remain constant over time. For example, facilities with managerial inefficiencies may provide poor quality of care and have longer wait times. The poorer quality of care also increases the risk of poorer health outcomes independent of wait times. Facility fixed effects eliminate any potential confounding from such permanent differences in quality across facilities.

We also include month fixed effects to control for seasonal variation in health outcomes and a time variable that counts up the number of observations for each person (e.g. time=1 in the first uncensored month, 2 in the second uncensored month). These controls mean that any estimated relationship between wait times and health outcomes is identified exclusively by within-facility variations over time and is independent of national seasonal variations or trends. After isolating the identifying variation in wait times to within-facility variations over time, significant variation in wait times remained with 10% of the facility-months having a wait time of less than 36 days and 10% having a wait time of more than 61 days (Table 1). Thirty-five percent of the wait time variation in our sample was attributed to within-facility variations over time (163/470=0.35).

Risk Adjustment and Other Explanatory Variables

We risk adjust models to control for observable differences in prior individual health status. Explanatory variables include age, gender, race, comorbidities, distance to nearest VA medical center and VA community based outpatient clinic, and VA priority status. Veterans with a service connected disability (priority status 1-3) receive priority access and likely experience shorter waits.[2]

We extract diagnosis codes from all inpatient and outpatient encounters financed by VA and Medicare during the 2001 baseline period. Our models include twenty-eight comorbidity indicator variables defined by Elixhauser et al. (1998) using ICD-9-CM diagnosis codes to risk adjust the outcomes. These comorbidities include a wide variety of physical and mental conditions such as heart disease, neurological disorders, depression, and substance abuse.

We use the diabetes severity index developed by Young et al. (2008) to control for diabetes severity. This index includes measures of complications from retinopathy, nephropathy, neuropathy, cerebrovascular disease, cardiovascular disease, peripheral vascular disease and metabolic disease. The individual is assigned a 0, 1, or 2 for each complication depending on its presence and severity (patients with neuropathy are given a 1 only; 6*2=12+1=13). ICD-9-CM codes from VA and Medicare outpatient and inpatient data for 2001 are used to determine the diabetes severity index for each individual.

Outcomes

We predict five outcomes: 1) monthly VA primary care utilization 2) mortality, 3) acute myocardial infarction (AMI), 4) stroke and 5) ACSC hospitalization (see Appendix B for a list of ACSC hospitalizations). We ascertain mortality using the VA Vital Status File, which determines the date of death from VA, Medicare and Social Security Administration data (Arnold et al. 2006). We use ICD-9-CM codes from VA and Medicare inpatient data to identify AMI, stroke and ACSC hospitalization. Methodology from Petersen et al. (1999) and Kiyota et al. (2004) defines AMIs, methodology from Rekker et al. (2002) defines strokes, and the AHRQ methodology (2010) determines ACSC hospitalizations. Following Prentice and Pizer (2008), hospitalizations are categorized as having one or more during the month versus none.

Analyses

*Model Set-up*

Data are analyzed using STATA 10.0. Our main explanatory variable of interest is VA primary care wait time. Patients in the hospital, nursing home or Medicare HMO during the wait time measurement period are not affected by VA outpatient waiting times, so we censor months if the veteran is institutionalized or in a Medicare HMO for all 3 months during the wait time measurement period. For AMI, stroke and ACSC hospitalization, we censor the outcome month if the individual was enrolled in a Medicare HMO during that month due to incomplete inpatient claims. Each individual has up to 21 observations. The outcome period is 24 months (2002 and 2003), but the lagged 3 month wait time excludes the first 3 months for each individual. The final sample size is 116,292 individuals.

We hypothesize that longer VA primary care waits will lead to less VA primary care utilization per veteran. To test this hypothesis we use a negative binomial model to predict the number of VA primary care visits in a month (Table 3A).[3] Coefficients from negative binomial models are difficult to interpret. Consequently, we run a policy simulation to illustrate the effect of increasing VA primary care wait times by one standard deviation (20.7 days) on VA primary care utilization (Table 3B). To examine the effect of wait times on health outcomes, we estimate logistic regression models that predict the probability of experiencing each outcome in each month of the outcome period (Table 4).[4]

Initial models include everyone in the sample. We hypothesize that longer waits for primary care will significantly increase the risk of poor health outcomes for the elderly. To disentangle the effect of age alone versus more severe diabetes or higher comorbidity burden we stratify the sample based on whether or not the patient was diagnosed during baseline with any of a set of "selected health conditions," which we chose because they are related to increased risk of the poor health outcomes included in the study. We define "selected health condition" as being diagnosed with any diabetes complication according to Young et al. (2008) or being diagnosed with congestive heart failure, cardiac arrhythmias, valvular disease, peripheral vascular disease, renal failure or obesity according to Elixhauser et al. (1998). We then stratify based on broad age groups and having a selected health condition at baseline. Finally, we stratify the analyses by five-year age groups since the risk of developing diabetes complications, common geriatric conditions and general frailty that may compromise diabetes management increases with age (Tables 3 and 4). Due to small sample sizes we could not stratify on both five-year age groups and the presence of a selected health condition.

**Results**

Like other samples of older veterans, our sample was predominantly male and in generally poor health (Table 2). Looking across the rows in Table 2 for the health status variables, the prevalence of each comorbidity increased with age, with a few exceptions such as paralysis and mental health disorders. These conditions were more prevalent in the less than 65 group, likely reflecting a population that was eligible for Medicare because of disability. Overall, 72% of the sample was considered to have one of the selected health conditions during the baseline period with this percentage ranging from 73 to 75% for veterans aged 70 or older. The percentages of individuals who died, experienced an AMI, stroke or ACSC hospitalization increased with the presence of a selected health condition at baseline and with age (Appendix C).

Veterans who visited VA facilities with longer primary care wait times had significantly lower primary care utilization compared to veterans who visited VA facilities with shorter primary care wait times for the entire sample and for each stratified group (Table 3A).[5] The size of the waiting time effect is illustrated in Table 3B, which predicts how VA primary care utilization would change with a one standard deviation increase (20.7 days) in wait time. For example, our model predicts that in the overall sample veterans had 0.349 primary care visits each month. An increase in wait times of 20.7 days was predicted to decrease the number of primary care visits to 0.338 visits per month. This translated into a predicted decrease in primary care utilization of 3.2. Generally, an increase of one standard deviation in wait time is predicted to decrease primary care utilization between 2 and 4%.

Longer waits for primary care did not significantly increase the odds of experiencing a poor health outcome for the overall sample (Table 4).[5] In analyses stratified by age or selected health conditions, wait time did not consistently increase the odds of experiencing a negative health outcome. Forty stratified models were run and wait time significantly increased the effect of experiencing a poor health outcome at $P<0.10$ in seven of these models. Except for the likelihood of experiencing an AMI, for which waiting had a significant effect at *P*=0.10 for veterans with a selected health condition at baseline, longer waits for primary care did not significantly predict poor health outcomes for veterans when the presence of a selected condition was considered but age was not. Longer waits for primary care significantly increased the risk of experiencing poor health outcomes for veterans aged 70 or more and diagnosed with a selected condition. A 10-day increase in facility-level primary care wait time led to a 2% (*P*=0.09) increase in the adjusted odds of mortality and a 6% increase in the adjusted odds of stroke (*P*=0.05) for this group. There was no significant relationship between primary care wait times and health outcomes for veterans aged less than 70 who were diagnosed with the selected conditions or veterans aged 70 or more who did not have these conditions. When stratifying by age only, the primary care wait time did not significantly increase the odds of experiencing a negative health outcome for veterans aged less than 70 years old. A 10-day increase in facility-level primary care wait time led to a 9% increase in the adjusted odds of stroke (*P* = 0.04) for veterans aged 70 to 74, and a 5% (*P* = 0.08) increase in the adjusted odds of AMI for veterans aged 75 to 79. For veterans 80 and older, the same 10-day facility-level change in wait time was associated with a 4% (*P*=0.04) increase in the adjusted odds of mortality and a 3% (*P*=0.04) increase in the adjusted odds of ACSC hospitalization.

**Discussion**

There is a strong consensus that complications from diabetes can be prevented or minimized if the disease is consistently managed in an outpatient setting (ADA 2005; Ross 2004; Stratton et al. 2000; DCCT 1996, 2000). Long waits for outpatient care are hypothesized to be problematic because decreases in utilization compromise the management of diabetes and increase the risk of poor health outcomes. In this sample of veterans with diabetes, longer waits for primary care modestly reduced primary care utilization (Table 3). An increase in wait times of 21 days is predicted to decrease primary care utilization by only 2 to 4%.

Results from this study provide no consistent support for the hypothesis that longer wait times negatively affect health outcomes, a general finding best illustrated by the lack of a significant relationship between longer primary care wait times and health outcomes for the overall sample (Table 4). Long wait times were hypothesized to significantly increase the risk of poor health outcomes, especially among elderly patients. Stratified analyses by selected health condition and age found small significant effects on health outcomes in only a few cases (Table 4).

Two aspects of our study design may cause our results to be overestimated. First, analyses include facility-level fixed effects, month dummies and a time trend. This design isolates the effect of wait times on health outcomes by focusing on month-to-month changes in wait times within facilities and by controlling for aspects of facility quality that remain constant over time. The principal limitation of this design is that we cannot completely rule out alternative explanations for our findings, including reverse causation and omitted variables. An example of reverse causation is an unobserved local flu epidemic at a VA facility that may increase wait times and cause poorer health outcomes that are not attributable to longer wait times. An example of omitted variable bias is a change in facility management that could simultaneously improve quality, reduce waiting times and improve outcomes through better resource allocation. The key to these explanations is that they feature unobservable local changes that are correlated with both local waiting time variations and local outcome changes. Either of these situations would cause our estimates to be overstated. In the absence of truly random variation in waiting times, this limitation is unavoidable. Second, it is also possible that our use of the number of days until the next available appointment for new patients as our wait time measure when our study population consists of established patients could cause our estimates to be overstated. This would occur if established patients do not experience longer waits when new patients do. We chose to use the new patient wait time because of its reliability (VHA 2008) and because our own experience suggests that congestion affects both new and established patients. These limitations in combination with an inconsistent pattern in the few significant results means we cannot formally conclude that wait times had a substantial effect on long-term health outcomes.

Our exogenous measure of wait time may also cause our results to be underestimated. By construction this variable is removed from the patient and is therefore an imprecise measure of congestion faced by individual patients. Imprecision in independent variables, or measurement error, causes estimated effects to be attenuated (biased towards zero). For this reason, the true effects of wait times on utilization and health outcomes could be larger than our estimates.

The lack of a significant effect of wait times on the long-term health outcomes in this study is divergent from previous research that finds a significant effect of timely access to care on short-term health outcomes in a population with diabetes. Subramanian et al. (2009) and Prentice et al. (2011) found patients visiting clinics providing more timely access had lower $HbA_{1c}$ levels compared to patients visiting clinics with less timely access. Quality improvement programs that achieve the largest improvements in $HbA_{1c}$ levels include timely access to care components, such as pharmacists who can adjust medication without waiting for physician authorization (Shojania et al. 2006). One reason why timely access might improve short-term outcomes is that a "white coat effect" has been found on medication adherence where patients are more likely to adhere to medications right before and after doctor's appointments (Cramer, Scheyer, and Mattson 1990; Ho, Bryson, and Rumsfeld 2009). Decreased overall utilization due to long waits as found in this study may decrease medication adherence, which in turn compromises $HbA_{1c}$, blood pressure and lipid control. The discrepancy between our findings and previous research predicting short-term outcomes may be simply due to the fact that long-term outcomes take a long time to develop whereas short-term health outcomes can respond more quickly to longer waits and decreased utilization.

Although we cannot conclude there was any significant effect on health outcomes in this study, findings from this research highlight the need for further study of the effect of wait times on long-term outcomes, especially among the elderly. The few significant effects found in this study corresponded to those over age 70 with one of the selected conditions or in certain narrower age groups and they all point in the same direction, with longer waits associated with undesirable outcomes (Table 4). This is consistent with previous studies that found relationships between waiting for outpatient care and mortality and ACSC hospitalizations for a sample of geriatric veterans (Prentice and Pizer 2007, 2008).

Despite the need for additional research, there are ongoing policy implications to which this study is relevant. Managers and policymakers debate the most effective scheduling policies to best serve the needs of all patients (Mehrotra, Keehl-Markowitz, and Ayanian 2008; Murray and Berwick 2003). Proponents of open access (e.g. next day (OA)) scheduling argue that the process of triaging decreases supply and unnecessarily increases demand (Murray and Berwick 2003). Others argue that the cost of triaging is smaller than the health benefits of providing priority access to patients with observable medical needs (Gravelle and Siciliani 2009).

VA official policy requires VA medical centers to provide care for urgent health care needs but then prioritizes patients' access to other care by the presence and extent of a service-connected disability (VHA 2002, 2006). Our results suggest this is a reasonable approach for patients diagnosed with diabetes. Veterans who visited VA facilities with longer waits had slightly less primary care utilization compared to veterans who visited facilities with shorter waits but there was no significant increase in the risk of experiencing a long-term health outcome for the former group.

The lack of a significant effect of waiting for health care on health outcome among this population of patients does not preclude a negative effect of waiting for care on other patient populations. Studies have repeatedly found that longer waits for outpatient care compromise short and long-term health outcomes on a variety of patient populations including patients with diabetes (Prentice et. al 2011; Subramanian et al. 2009) and vulnerable geriatric clinic populations (Prentice and Pizer 2007; 2008). In contrast, Subramanian et al. (2009) found that diabetes processes of care (e.g. HbA1c testing) suffered in OA clinics versus control clinics resulting in fears that OA scheduling increased focus on acute care management and crowded out long-term chronic disease management. If so, getting patients into outpatient care as soon as possible may not be the most effective approach and having everyone wait longer to ensure appropriate processes are followed may be better. Policymakers and clinicians should think carefully about the cost and benefits of scheduling procedures for different patient populations to ensure high-quality care and prevent poor health outcomes.

**Notes**

1. For ease of presentation, "facility" and "parent station" are used interchangeably throughout the paper to refer to a VA medical center.

2. Service-connected disability is a condition that the VA has determined was incurred in or aggravated by military service.

3. We also test the hypothesis that veterans would substitute Medicare for VA care and total utilization would not go down. We estimate negative binomial models predicting total Medicare outpatient utilization using average VA wait time and find no significant relationship.

4. Sample sizes for each negative binomial and logistic regression model are included in Tables 3 and 4. Sample sizes change for each model because different months are censored due to hospitalization, nursing home stays and Medicare HMO enrollment. Finally, if there was no variation in outcome for all observations under a VA facility or for all individuals diagnosed with the same comorbidity all observations from that VA facility or diagnosed with the comorbidity are dropped. This was more common in the AMI and stroke models where the outcomes are rare.

5. Full results from utilization and health outcome models are available from the authors upon request.

## References


Agency for Health Care Research and Quality (AHRQ). 2001. "AHRQ Quality Indicators-Guide to Prevention Quality Indicators: Hospital Admission for Ambulatory Care Sensitive Conditions". Rockville, MD.

Agency for Health Care Research and Quality (AHRQ). 2010. http://www.qualityindicators.ahrq.gov/pqi_download.htm; Accessed on March 30, 2010.

American Diabetes Association (ADA), 2005. "Standards of Medical Care in Diabetes." *Diabetes Care* 28(1):S4-S36.

Arnold, N., M.-W. Sohn, C. Maynard, and D. M. Hynes. 2006. "VA-NDI Mortality Merge Project". *VIReC Technical Report 2*. pp. 1-34. Hines, IL: VA Information Resource Center.

Baar, B. 2005a. "New Patient Monitor: Data Definitions". Veterans Health Administration Support Services Center.

Baar, B. 2005b. "Performance Measure Clinic Data Definitions and Discussion". Veterans Health Administration Support Services Center.

Buchmueller, T. C., K. Grumbach, R. Kronick, and J. G. Kahn. 2005. "The Effect of Health Insurance on Medical Care Utilization and Implications for Insurance Expansion: A Review of the Literature." *Medical Care Research and Review* 62(1):3-30.

Chau, D. and S. V. Edelman. 2001. "Clinical Management of Diabetes in the Elderly." *Clinical Diabetes* 19(4):172-74.

Cramer, J. A., R. D. Scheyer, and R. H. Mattson. 1990. "Compliance Declines Between Clinic Visits." *Archives of Internal Medicine* 150:1509-10.

Congressional Budget Office (CBO), 2008. "Key Issues in Analyzing Major Health Insurance Proposals". pp. 1-167. Washington D.C.: Congressional Budget Office.

Cunningham, P. J. and L. E. Felland. 2008. "Falling Behind: Americans' Access to Medical Care Detoriates, 2003-2007". Washington D.C.: Center for Studying Health System Change.

Diabetes Control and Complications Trial (DCCT). 1996. "Lifetime Benefits and Costs of Intensive Therapy as Practiced in the Diabetes Control and Complications Trial." *Journal of the American Medical Association* 276:1409-15.

Diabetes Control and Complications Trial (DCCT). 2000. "Retinopathy and Nephropathy in Patients with Type 1 Diabetes Four Years After a Trial of Intensive Insulin Therapy." *New England Journal of Medicine* 342:381-89.

Durso, S. C. 2006. "Using Clinical Guidelines Designed for Older Adults With Diabetes Mellitus and Complex Health Status." *Journal of the American Medical Association* 295(16):1935-40.

Elixhauser, A., C. Steiner, D. R. Harris, and R. M. Coffey. 1998. “Comorbidity Measures for Use with Administrative Data.” *Medical Care* 36(1):8-27.

Fried, L.P., L. Ferrucii, J. Darer, J.D. Williamson and G. Anderson. 2004. "Untangling the Concepts of Diasbilty, Frailty, and Comorbidity: Implications for Improved Targeting and Care." *Journal of Gerontology* 59(3):255-263.

Fowler, M. J. 2008. "Microvascular and Macrovascular Complications of Diabetes" *Clinical Diabetes* 26(2): 77-82.

Gravelle, H. and L. Siciliani. 2009. “Third Degree Waiting Time Discrimination: Optimal Allocation of a Public Sector Healthcare Treatment Under Rationing by Waiting.” *Health Economics* 18:977-86.

Helmer, D. A., U. Sambamoorthi, M. Rajan, C-L. Tseng, and L. M. Pogach. 2008. "Glycemic Control in Elderly Veterans with Diabetes: Individualized, Not Age-Based." *Diabetes Care* 31(4):728-731.

Ho, P. M., C. L. Bryson, and J. S. Rumsfeld. 2009. “Medication Adherence: Its Importance in Cardiovascular Outcomes.” *Circulation* 119:3028-35.

Hornick, T. and D. C. Aron. 2008. “Managing Diabetes in the Elderly: Go easy, Individualize.” *Cleveland Clinic Journal of Medicine* 75(1):70-78.

Institute of Medicine (IOM). The Committee on the Quality of Health Care in America. 2001. “Crossing the Quality Chasm: A New Health System for the 21st Century”. Washington D.C.: Institute of Medicine

Kenagy, J. W., D. M. Berwick, and M. F. Shore. 1999. “Service Quality in Health Care.” *Journal of the American Medical Association* 281(7):661-65.

Kiyota, Y., S. Schneeweiss, R. J. Glynn, C. C. Cannuscio, J. Avorn, and D. H. Solomon. 2004. “Accuracy of Medicare Claims-Based Diagnosis of Acute Myocardial Infarction: Estimating Positive Predictive Value on the Basis of Review of Hospital Records.” *American Heart Journal* 148(1):100-04.

Mayfield, J. A., G. E. Reiber, L. J. Sanders, D. Jannisse, and L. M. Pogach. 1998. “Preventive Foot Care in People With Diabetes.” *Diabetes Care* 21(12):2161-77.

Mehrotra, A., L. Keehl-Markowitz, and J. Z. Ayanian. 2008. “Implementing Open-Access Scheduling of Visits in Primary Care Practices: A Cautionary Tale.” *Annals of Internal Medicine* 148(12):915-22.

Miller, D.R., and L.M. Pogach. 2008. "Longitudinal Approaches to Evaluate Health Care Quality and Outcomes: The Veterans Health Administration Diabetes Epidemiology Cohort" *J Diabetes Sci Technol* 2(1): 24-32.

Murray, M. and D. M. Berwick. 2003. “Advanced Access: Reducing Waiting and Delays in Primary Care.” *Journal of the American Medical Association* 289:1035-40.

Petersen, L. A., S. Wright, S.-L. T. Normand, and J. Daley. 1999. “Positive Predictive Value of the Diagnosis of Acute Myocardial Infarction in an Administrative Database.” *Journal of General Internal Medicine* 14:555-58.

Prentice, J. C., B. G. Fincke, D. R. Miller, and S. D. Pizer. 2011. “Outpatient Waiting Times and Diabetes Care Quality Improvement.” *American Journal of Managed Care* 17(2):e43-e54.

Prentice, J. C. and S. D. Pizer. 2007. “Delayed Access to Health Care and Mortality.” *Health Services Research* 42(2):644-62.

Prentice, J. C. and S. D. Pizer. 2008. “Waiting Times and Hospitalizations for Ambulatory Care Sensitive Conditions.” *Health Services Outcomes and Research Methodology* 8:1-18.

Reker, D. M., A. K. Rosen, H. Hoenig, D. R. Berlowitz, J. Laughlin, L. Anderson, C. R. Marshall, and M. Rittman. 2002. “The Hazards of Stroke Case Selection Using Administrative Data.” *Medical Care* 40(2):96-104.

Ross, S. A. 2004. “Controlling Diabetes: The Need for Intensive Therapy and Barriers in Clinical Management.” *Diabetes Research and Clinical Practice* 65S:S29-S34.

Shojania, K. G., S. R. Ranji, K. M. McDonald, J. M. Grimshaw, V. Sundaram, R. J. Rushakoff, and D. K. Owen. 2006. “Effects of Quality Improvement Strategies for Type 2 Diabetes on Glycemic Control: A Meta-Regression Analysis.” *Journal of the American Medical Association* 296:427-40.

Stratton, I. M., A. I. Adler, H. A. W. Neil, D. R. Matthews, S. E. Manley, C. A. Cull, D. Hadden, R. C. Turner, and R. R. Holman. 2000. “Association of Glycamia With Macrovascular and Microvascular Complications of Type 2 Diabetes (UKPDS 35): Prospective Observational Study.” *British Medical Journal* 321(7258):405-12.

Subramanian, U., R. T. Ackermann, E. J. Brizendine, C. Saha, M. B. Rosenman, D. R. Willis, and D. G. Marrero. 2009. “Effect of Advanced Access Scheduling on Processes and Intermediate Outcomes of Diabetes Care and Utilization.” *Journal of General Internal Medicine* 24(3):327-33.

United States General Accounting Office (U.S. GAO). 1993. “VA Health Care: Restructuring Ambulatory Care System Would Improve Services to Veterans”. Washington D.C.: United States General Accounting Office.

Van Deusen Lukas, C., M. Meterko, D. Mohr, and M. N. Seibert. 2004. “The Implementation and Effectiveness of Advanced Clinic Access”. Health Services Research and Development Management Decision and Research Center, Boston: Office of Research and Development, Department of Veterans Affairs.

Veterans Health Administration (VHA) 2008. “PMO- New and Established Patients”. Department of Veterans Affairs.

Veterans Health Administration (VHA) 2002. “Priority for Outpatient Medical Services and Inpatient Hospital Care”. Department of Veterans Affairs.

Veterans Health Administration (VHA) 2006. "Process for Ensuring Timely Access to Outpatient Clinical Care”. Department of Veterans Affairs.

Young, B. A., E. Lin, M. V. Korff, G. Simon, P. Ciechanowski, E. J. Ludman, S. Everson-Stewart, L. Kinder, M. Oliver, E. J. Boyko, and W. J. Katon. 2008. “Diabetes Complications Severity Index and Risk of Mortality, Hospitalization, and Healthcare Utilization.” *American Journal of Managed Care* 14(1):15-24.

**Table 1: Descriptive statistics of Unadjusted and Adjusted Lagged 3 Month Average Wait Time (n=2,541)**

| Distribution | Unadjusted Wait Time | Adjusted Wait Time* |
|---|---|---|
| 10% | 25.79 | 35.66 |
| 25% | 34.40 | 41.97 |
| 50% | 45.44 | 48.14 |
| 75% | 57.67 | 53.99 |
| 90% | 72.96 | 61.29 |
| Mean | 48.44 | 48.44 |
| Variance | 470.30 | 163.13 |

*Adjusted for facility-level fixed effects, month dummies and time since entering the sample.

**Table 2: Descriptive Demographic and Comorbidity Statistics Stratified by Age**

| | <65 years old n=36,088 | 65-69 years old n=27,042 | 70-74 years old n=22,790 | 75-80 years old n=18,946 | >=80 years old n=11,426 | Total sample n=116,292 |
|---|---|---|---|---|---|---|
| Male (%) | 97.5 | 98.6 | 99.1 | 98.0 | 97.7 | 98.2 |
| White (%) | 71.6 | 76.1 | 73.1 | 74.6 | 73.2 | 73.6 |
| Black (%) | 23.0 | 20.4 | 21.4 | 20.0 | 21.5 | 21.5 |
| Hispanic (%) | 2.4 | 1.4 | 3.7 | 3.8 | 3.5 | 2.8 |
| Asian (%) | 0.5 | 0.4 | 0.4 | 0.5 | 0.6 | 0.5 |
| Native American (%) | 0.9 | 0.5 | 0.5 | 0.4 | 0.3 | 0.6 |
| Other (%) | 1.7 | 1.2 | 0.9 | 0.6 | 0.9 | 1.2 |
| Priority status 1-3 (%) | 53.6 | 28.6 | 29.9 | 35.7 | 41.1 | 39.0 |
| | | | | | | |
| Diabetes Severity Index (mean, S.D.) | 1.5 (1.7) | 1.6 (1.7) | 1.7 (1.8) | 1.8 (1.8) | 1.8 (1.8) | 1.6 (1.7) |
| Congestive heart failure (%) | 9.1 | 9.5 | 11.2 | 13.3 | 15.3 | 10.9 |
| Cardiac arrhythmias (%) | 5.9 | 9.0 | 10.7 | 13.8 | 16.9 | 9.9 |
| Valvular disease (%) | 1.8 | 2.7 | 3.3 | 4.4 | 4.7 | 3.0 |
| Pulmonary circulation disorder (%) | 0.6 | 0.5 | 0.5 | 0.6 | 0.7 | 0.6 |
| Peripheral vascular disease (%) | 9.7 | 12.4 | 13.9 | 16.0 | 15.7 | 12.8 |
| Hypertension (%) | 69.3 | 77.3 | 77.8 | 77.8 | 75.1 | 74.8 |
| Paralysis (%) | 2.0 | 1.0 | 1.2 | 1.2 | 1.1 | 1.4 |
| Other neurological disorder (%) | 3.6 | 2.3 | 2.8 | 3.5 | 3.9 | 3.1 |
| Chronic pulmonary disease (%) | 15.0 | 14.9 | 16.8 | 17.9 | 16.9 | 16.0 |
| Hypothyroidism (%) | 4.4 | 4.1 | 4.8 | 5.5 | 6.6 | 4.8 |
| Renal failure (%) | 6.0 | 5.3 | 6.1 | 7.2 | 7.7 | 6.2 |
| Liver disease (%) | 4.0 | 1.4 | 1.2 | 0.8 | 0.6 | 2.0 |
| Peptic ulcer disease including bleeding (%) | 1.5 | 1.9 | 2.0 | 2.3 | 2.2 | 1.9 |
| AIDS (%) | 0.7 | 0.2 | 0.1 | 0.03 | 0.04 | 0.3 |
| Lymphoma (%) | 0.5 | 0.6 | 0.7 | 0.7 | 0.7 | 0.6 |
| Metastatic Cancer (%) | 0.4 | 0.5 | 0.7 | 0.7 | 0.8 | 0.6 |
| Solid Tumor Without Metastasis (%) | 5.5 | 11.2 | 14.5 | 17.9 | 19.8 | 12.0 |
| Rheumatoid arthritis /Collagen vascular diseases (%) | 2.1 | 1.7 | 1.8 | 1.6 | 1.9 | 1.8 |
| Coagulopathy (%) | 1.6 | 1.6 | 1.6 | 1.9 | 2.1 | 1.7 |
| Obese (%) | 22.9 | 17.6 | 14.1 | 11.0 | 6.5 | 16.4 |
| Weight loss (%) | 1.1 | 1.1 | 1.4 | 1.9 | 2.4 | 1.4 |
| Fluid and electrolyte disorders (%) | 4.3 | 3.9 | 4.3 | 4.6 | 6.0 | 4.4 |

| | | | | | | |
|---|---|---|---|---|---|---|
| Blood loss anemia (%) | 0.2 | 0.3 | 0.5 | 0.6 | 0.5 | 3.8 |
| Deficiency anemias (%) | 6.5 | 7.3 | 8.8 | 11.4 | 14.0 | 8.7 |
| Alcohol abuse (%) | 6.4 | 2.4 | 1.8 | 1.2 | 0.8 | 3.2 |
| Drug abuse (%) | 3.9 | 0.5 | 0.3 | 0.2 | 0.1 | 1.4 |
| Psychoses (%) | 21.6 | 6.1 | 5.5 | 5.6 | 5.4 | 10.6 |
| Depression (%) | 17.6 | 7.7 | 7.2 | 7.3 | 7.1 | 10.5 |
| Number of comorbidities (mean, S.D.) | 2.3 (1.6) | 2.0 (1.5) | 2.2 (1.6) | 2.3 (1.6) | 2.4 (1.7) | 2.2 (1.6) |
| Selected Health Condition† | 71.1 | 71.4 | 73.1 | 74.7 | 73.4 | 72.4 |

†This is defined as anyone diagnosed with retinopathy, neuropathy, nephropathy, cardiovascular, cerebrovascular, peripheral vascular or metabolic disease according to Young et al.(2008) or anyone with congestive heart failure, cardiac arrhythmias, valvular disease, peripheral vascular disease, renal failure or obesity according to Elixhauser et al. (1998).

**Table 3A: Coefficient, Standard Errors and P-Values of Linear 3 month Average Wait time Predicting VA Primary Care Utilization Stratified by Age and Selected Health Condition (SHC)†**

| Stratification | **β*** | **Standard Error** | ***P*-Value** |
|---|---|---|---|
| None (n=116,113) | -0.0150 | 0.0011 | <0.0001 |
| No SHC (n=32,080) | -0.0173 | 0.0022 | <0.0001 |
| At least one SHC (n=84,033) | -0.0142 | 0.0013 | <0.0001 |
| <70 and no SHC (n=18,137) | -0.0178 | 0.0031 | <0.0001 |
| < 70 and SHC (n=44,881) | -0.0112 | 0.0019 | <0.0001 |
| >=70 and no SHC (n=15,463) | -0.0164 | 0.0032 | <0.0001 |
| >70 and SHC (n=43,116) | -0.0174 | 0.0018 | <0.0001 |
| <65 years old (n=36,016) | -0.0110 | 0.0022 | <0.0001 |
| 65-69 years old (n=33,095) | -0.0149 | 0.0026 | <0.0001 |
| 70-74 years old (n=28,236) | -0.0173 | 0.0025 | <0.0001 |
| 75-79 years old (n=22,704) | -0.0159 | 0.0028 | <0.0001 |
| >=80 years old (n=14,313) | -0.0189 | 0.0036 | <0.0001 |

**Table 3B: Policy Simulation Predicting Change in Monthly VA Primary Care Utilization With A One Standard Deviation Increase in Wait Times Stratified by Age and Selected Health Condition (SHC)†**

| | **Baseline** | **Simulation** | **% Change** |
|---|---|---|---|
| Average wait time | 47.9 | 68.6 | 43.2 |
| Stratification | | | |
| None (n=116,113) | 0.349 | 0.338 | -3.2 |
| No SHC (n=32,080) | 0.296 | 0.286 | -3.4 |
| At least one SHC (n=84,033) | 0.369 | 0.359 | -2.7 |
| <70 and no SHC (n=18,137) | 0.291 | 0.280 | -3.8 |
| < 70 and SHC (n=44,881) | 0.358 | 0.350 | -2.2 |
| >=70 and no SHC (n=15,463) | 0.303 | 0.293 | -3.7 |
| >70 and SHC (n=43,116) | 0.381 | 0.368 | -3.4 |
| <65 years old (n=36,016) | 0.336 | 0.328 | -2.4 |
| 65-69 years old (n=33,095) | 0.341 | 0.331 | -2.9 |
| 70-74 years old (n=28,236) | 0.356 | 0.343 | -3.7 |
| 75-79 years old (n=22,704) | 0.365 | 0.353 | -3.3 |
| >=80 years old (n=14,313) | 0.361 | 0.347 | -3.9 |

†Selected health condition is defined as anyone diagnosed with retinopathy, neuropathy, nephropathy, cardiovascular, cerebrovascular, peripheral vascular or metabolic disease according to Young et al.(2008) or anyone with congestive heart failure, cardiac arrhythmias, valvular disease, peripheral vascular disease, renal failure or obesity according to Elixhauser et al. (1998).

*The coefficient, standard error and *P*-value for the VA wait time is shown. Wait time is divided by 10 so the coefficient reflects a 10 day increase in wait time. Models also include Elixhauser comorbidities, diabetes severity index, distance to nearest VA medical center, distance to nearest VA outpatient clinic, station-level fixed effects, demographics, time since entering the sample and month dummies. Full results are available from the authors upon request.

**Table 4: Effects of Linear 3-month Average Wait Time on Outcomes Stratified by Age and Selected Health Condition (SHC)†**

| | Mortality | | | AMI | | | Stroke | | | ACSC hospitalization | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Stratification | n‡ | A.O.R.^ | *P*-value* | n | A.O.R | *P*-value | n | A.O.R. | *P*-value | n | A.O.R. | *P*-value |
| None | 116,292 | 1.01 | 0.52 | 116,113 | 1.02 | 0.31 | 116,113 | 1.01 | 0.47 | 116,113 | 1.00 | 0.88 |
| No SHC | 32,129 | 0.99 | 0.57 | 30,640 | 0.94 | 0.14 | 31,230 | 1.00 | 0.99 | 32,080 | 1.00 | 0.85 |
| At least one SHC | 84,163 | 1.01 | 0.33 | 84,033 | **1.03** | **0.10** | 83,869 | 1.02 | 0.44 | 84,033 | 1.00 | 0.78 |
| < 70 and SHC° | 44,956 | 0.99 | 0.51 | 44,881 | 1.02 | 0.51 | 44,178 | 0.95 | 0.16 | 44,881 | 0.99 | 0.40 |
| >=70 and no SHC | 15,469 | 0.99 | 0.71 | 13,405 | 0.95 | 0.28 | 13,594 | 0.97 | 0.61 | 15,415 | 1.03 | 0.21 |
| >=70 and SHC | 43,177 | **1.02** | **0.09** | 42,952 | 1.04 | 0.12 | 42,780 | **1.06** | **0.05** | 43,116 | 1.00 | 0.71 |
| <65 years old | 36,081 | 1.01 | 0.76 | 35,112 | 1.04 | 0.28 | 33,849 | 0.96 | 0.29 | 36,016 | 1.00 | 0.81 |
| 65-69 years old | 33,184 | 0.98 | 0.26 | 32,848 | 0.98 | 0.52 | 31,368 | 0.98 | 0.70 | 33,095 | 0.98 | 0.20 |
| 70-74 years old | 28,282 | 1.00 | 0.95 | 27,777 | 0.99 | 0.85 | 27,358 | **1.09** | **0.04** | 28,164 | 0.99 | 0.43 |
| 75-79 years old | 22,709 | 1.01 | 0.76 | 22,330 | **1.05** | **0.08** | 21,811 | 1.01 | 0.78 | 22,704 | 1.00 | 0.77 |
| >=80 years old | 14,328 | **1.04** | **0.04** | 13,823 | 1.00 | 0.95 | 13,328 | 1.03 | 0.52 | 14,313 | **1.03** | **0.04** |

†Selected health condition is defined as anyone diagnosed with retinopathy, neuropathy, nephropathy, cardiovascular, cerebrovascular, peripheral vascular or metabolic disease according to Young et al.(2008) or anyone with congestive heart failure, cardiac arrhythmias, valvular disease, peripheral vascular disease, renal failure or obesity according to Elixhauser et al. (1998).

‡ n=number of persons in model. The sample size is different for each model due to different censoring restrictions (see Methods: Model Set Up and note 4 for a detailed description).

^The adjusted odds ratio for the VA wait time is shown. Wait time is divided by 10 so the adjusted odds ratio reflects a 10 day increase in wait time. Models also include distance to nearest VA medical center, distance to nearest VA outpatient clinic, station-level fixed effects, demographics, time since entering the sample and month dummies. Full results are available from the authors upon request.

* Results with P-values less than or equal to 0.10 shown in bold.

°The infrequency of outcomes in the <70 years old and no selected health condition group resulted in unstable models when using facility-fixed effects.

**Appendix A: Example of Exogenous Wait Time Calculation***

| | **2001** (Sample Selection Period) | | | **January 2002 Wait Times** | | **February 2002 Wait Times** | |
|---|---|---|---|---|---|---|---|
| Appointment Type | *Station 1* | *Station 2* | | *Station 1* | *Station 2* | *Station 1* | *Station 2* |
| | # of appointments | # of appointments | National proportion of appointments | Wait in days | Wait in days | Wait in days | Wait in days |
| Primary Care | 55 | 42 | 0.97 | 20.8 | 36.5 | 15.2 | 34.2 |
| Geriatric Primary Care | 1 | 1 | 0.02 | 33.2 | 28.6 | Missing | 10.0 |
| Women's Clinic | 1 | ---- | 0.01 | 21.2 | ---- | 22.5 | ---- |
| | | | Wait Time Calculation | (0.97*20.8)+ (0.02*33.2)+ (0.01*21.2)= **21.05** | (0.975*36.5)+ (0.025*28.6)= **36.30** | (0.97*15.2)+ (0.02*0)+ (0.01*22.5)= **14.97** | (0.975*34.2)+ (0.025*10.0)= **33.60** |

*There are three primary care appointment types: 1) Primary care, 2) Geriatric Primary Care and 3) Women's Clinic. Suppose there are only two parent stations. The total number of appointments during the sample selection period is 100. Overall, 97% of the appointments are in Primary Care, 2% were in Geriatric Primary Care, and 1% are in the Women's Clinic, even though these proportions actually differ by individual station (e.g. the proportions are 96% for Primary Care, 2% for Geriatric Primary Care and 2% for the Women's Clinic for station 1).

The wait time for station 1 in January is: (0.97*20.8) + (0.02*33.2)+(0.01*21.2)=21.05 days. Since this parent station uses all clinic appointment types and has no missing wait times, the wait time is constructed by multiplying the wait in days for each appointment type by the national proportion of appointments associated with each clinic type.

The wait time for station 1 in February is: (0.97*15.2)+(0.02*0.0)+(0.01*22.5)=14.97 days. For geriatric primary care appointments, the wait time is imputed with 0 because the wait time for this appointment type is missing in the February data. This indicates no next available appointments are scheduled for geriatric primary care in February. Since there is a wait time for geriatric primary care in January, it is assumed that there is no waiting for geriatric primary care appointments in February and patients could get in right away.

There are no appointments or wait times for women's clinic in either January or February for the Women's Clinic for Station 2. Station 2 is assumed not to use Women's clinic and the 1% of appointments that are attributed nationally to the Women's Clinic is redistributed evenly among the other 2 primary care appointment types so the proportions add up to 1. The proportions for Station 2 for primary care are 0.975 (compared to 0.97 nationally) and for geriatric primary care 0.025 (compared to 0.02 nationally).

All patients that lived closest to the same VA facility are assigned the same exogenous facility-month wait time corresponding to that facility.

**Appendix B: Descriptive Statistics of ACSC Hospitalizations Stratified by Age**

| Percent with each ACSC hospitalization | <65 years old n=36245 | 65-69 years old n=27155 | 70-74 years old n=22924 | 75-80 years old n=19118 | >=80 years old n=11588 | Total sample n=117030 |
|---|---|---|---|---|---|---|
| Bacterial pneumonia | 2.1 | 2.3 | 3.0 | 3.9 | 5.8 | 3.0 |
| Dehydration | 0.7 | 0.8 | 0.9 | 1.5 | 2.2 | 1.0 |
| Urinary tract infection | 0.8 | 1.0 | 1.6 | 2.2 | 3.6 | 1.5 |
| Perforated appendix | 0.07 | 0.06 | 0.06 | 0.06 | 0.04 | 0.06 |
| Angina | 0.7 | 0.6 | 0.7 | 0.9 | 0.8 | 0.7 |
| Congestive heart failure | 3.7 | 4.2 | 5.2 | 6.0 | 7.0 | 4.8 |
| Hypertension | 0.4 | 0.2 | 0.3 | 0.3 | 0.4 | 0.3 |
| Asthma | 0.2 | 0.2 | 0.2 | 0.1 | 0.2 | 0.2 |
| Chronic obstructive pulmonary disease | 1.2 | 1.4 | 1.9 | 1.9 | 2.1 | 1.6 |
| Uncontrolled diabetes | 0.5 | 0.3 | 0.3 | 0.4 | 0.4 | 0.4 |
| Diabetes-sort term complications | 0.8 | 0.4 | 0.3 | 0.4 | 0.4 | 0.5 |
| Diabetes-long term complications | 3.6 | 2.9 | 3.2 | 3.8 | 4.0 | 3.4 |
| Lower extremity amputations among diabetics | 1.6 | 1.4 | 1.4 | 1.3 | 1.3 | 1.4 |

**Appendix C: Descriptive Statistics of Outcomes Stratified by Age and Selected Health Condition (SHC)†**

| Stratification | Died (%) | Heart attack (AMI) (%) | Stroke (%) | ACSC Hospitalization (%) |
|---|---|---|---|---|
| None( n=116,292) | 10.2 | 2.8 | 2.0 | 15.0 |
| No SHC† (n=32,129) | 7.2 | 1.4 | 1.5 | 8.5 |
| At least one SHC (n=84,163) | 11.3 | 3.4 | 2.2 | 17.5 |
| Age<70 & no SHC (n=18,169) | 4.5 | 1.0 | 1.2 | 6.9 |
| Age<70 & SHC (n=44,961) | 7.8 | 2.7 | 1.8 | 15.1 |
| Age>=70 & no SHC (n=13,960) | 10.7 | 2.0 | 2.0 | 10.6 |
| Age>=70 & SHC (n=39,202) | 15.2 | 4.1 | 2.7 | 20.3 |
| Age<65years old (n=36,088) | 6.1 | 2.0 | 1.4 | 12.9 |
| Age 65-69 years old (n=27,042) | 7.9 | 2.6 | 1.9 | 12.5 |
| Age 70-74 years old (n=22,790) | 10.7 | 3.1 | 2.2 | 15.0 |
| Age 75-79 years old (n=18,946) | 14.0 | 3.7 | 2.6 | 18.3 |
| Age>=80 years old (n=11,426) | 20.7 | 4.3 | 2.9 | 22.3 |

†Selected health condition is defined as anyone diagnosed with retinopathy, neuropathy, nephropathy, cardiovascular, cerebrovascular, peripheral vascular or metabolic disease according to Young et al.(2008) or anyone with congestive heart failure, cardiac arrhythmias, valvular disease, peripheral vascular disease, renal failure or obesity according to Elixhauser et al. (1998).